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ABSTRACT

The Consequences of 'In-Work' Benefit Reform in Britain: New Evidence from Panel Data*

In October 1999, the British government enacted the Working Families' Tax Credit, a generous tax credit aimed at encouraging work among low-income families with children. This paper uses longitudinal data collected between 1991 and 2001 to evaluate the effect of this reform on single mothers. We identify this impact by comparing changes in behavior of lone mothers to changes for single women without children. Our results show that the financial incentives of the reform had powerful effects on a wide range of lone mothers' decisions. The reform led to a substantial increase in employment rates of about 7 percentage points, which was driven by both higher rates at which lone mothers remained in the labor force and higher rates at which they entered it. Women's responses were highly heterogeneous, with larger effects for mothers with one pre-school aged child, and virtually no effect for mothers with multiple older children. The reform also led to significant reductions in single mothers' subsequent fertility and in the rate at which they married. Our findings suggest that the generous childcare tax credit component of the reform played a key role in explaining the estimated employment responses. Finally, we find relatively large behavioral effects in anticipation of the actual reform, which emphasizes the importance of allowing for such effects in future evaluation research.

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

Keywords: difference-in-difference and panel-data estimators, welfare-to-work program evaluation, lone mothers, anticipation effects

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Over the past three decades many countries have witnessed a rapid growth in the proportion of households headed by single women with dependent children.[1] This increase was accompanied by low or declining employment rates of lone mothers relative to other women (OECD, 1998 and 2001), which in turn contributed to the dramatic increase in the proportion of children growing up poor (Cornia, 1997; UNICEF, 2000; Micklewright, 2003). In the United States, for instance, the poverty rate of children rose from about 15 percent in 1970 to about 23 percent in 1992 (Gottschalk and Danziger, 2001). Over the same period, child poverty rates in the United Kingdom increased from about 10 to 27 percent (Hill and Jenkins, 2001). Bradbury and Jäntti (2001) document similar trends for a number of other countries, including the Netherlands, Germany, Italy, Australia, and Russia.[2] Magnified by political frustration with the work disincentives imbedded in earlier welfare programs, these trends prompted several governments to change their traditional public assistance schemes, introducing or expanding welfare-to-work programs, reducing out-of-work benefit levels and rates, altering funding methods and administration, and encouraging marriage.[3]

In the middle of the 1980s the United States preceded other countries in this wave of reforming the welfare system by increasing the generosity of the Earned Income Tax Credit (EITC). Since then and through a series of successive expansions, EITC has emerged as a popular alternative method for transferring income to low-income families with children.[4] In addition, 1996 saw major welfare reform legislation replacing the federal Aid to Families with Dependent Children (AFDC) — the largest cash assistance program for low-income families — with the Temporary Assistance for Needy Families (TANF) block grant.[5] Some of these innovations were

[1] For a documentation of these trends in European countries and in the United States, see Eurostat (2002) and U.S. Census Bureau (2002) respectively.
[2] Concern among policy makers and the public about the growing proportion of household headed by lone mothers and their greater exposure to poverty experience was not without foundation. Several studies show that growing up in a nonintact family has detrimental consequences for children's wellbeing across a wide range of outcomes (e.g., McLanahan and Sandefur, 1994; Haveman and Wolfe, 1995; Duncan and Brooks-Gunn, 1997).
[3] For detailed descriptions of the major changes in welfare programs over the 1990s, especially in the United States and the United Kingdom, see Moffitt (1998 and 2003a), Card and Blank (2000), Blank (2002), and Blundell *et al*. (forthcoming).
[4] Blundell and Hoynes (forthcoming) discuss the contrasts between the EITC and traditional welfare benefits. On this issue and for further descriptions of the program, see also Scholz (1996), Blundell (2001), and Hotz and Scholz (2003). Evaluations of the effects of EITC on a wide range of outcomes are in Eissa and Liebman (1996), and Eissa and Hoynes (2000 and forthcoming).
[5] For further discussion on the transition from AFDC to TANF, see Moffitt (2003a).

subsequently adopted by a number of other industrialized countries, including the United Kingdom, which adjusted such welfare reforms to their specific institutional structures and economic conditions.[6] In October 1999, the United Kingdom replaced its version of the EITC, called Family Credit (FC), by enacting the Working Families' Tax Credit (WFTC) program, a more generous tax credit designed to improve the work incentives for families with low incomes. In May 2000 under the WFTC regime, a total of 548,000 lone-parent households received the tax credit, some 120,000 (almost 30 percent) more than those who received it in August 1999 under the FC regime (Inland Revenue, 2003). During this same period, the proportion of lone-parent families remained fairly stable, while single-mothers' employment rates increased by nearly 8 percentage points in just twenty-four months from about 41 percent at the end of 1997 to about 49 percent in early 2000 (Office for National Statistics, various years). In subsequent years, lone-mothers' employment rates stabilized at about 50 percent, while the number of lone-parent families in receipt of WFTC kept increasing, reaching 737,000 households by November 2002. Interestingly, these welfare and employment changes were accompanied by a significant drop in child poverty rates. For example, Brewer *et al*. (2003a) estimate that by 2001 the number of children in low-income households fell by more than half a million from about 4.4 million in 1996/7, an 11 percent reduction.

It is important to note that several other labor market reforms took place during this period, some of which may have had a direct impact on low-income single mothers. These include the introduction of various welfare-to-work programs, national minimum wage legislation, and changes to National Insurance contributions and other means-tested benefits. Section I will place WFTC into the context of such policy reforms and describe them in detail. The aim of this study, however, is to focus on what many analysts consider to be the dominant reform of that period, the introduction of WFTC. We therefore assess how much of the observed changes in lone parents' employment and other measurable behavior can be attributed to this reform.

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

[6] See Blundell (2001 and 2002), Brewer (2001) and Blundell and Hoynes (forthcoming) for a detailed account of the in-work reforms introduced in the United States, the United Kingdom and Canada over the 1990s.

parents, since it increased the financial payoffs to working any given hours level above 16 hours per week (Blundell *et al*., 2000).[7] But because of the interaction of WFTC with other benefits (see the discussion in Section I), greater income increases were expected to be observed 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 a static behavioral model of household labor supply with 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 (forthcoming) suggest that the employment impact of WFTC among lone mothers was indeed positive and modest. However, more recent results reported in Brewer *et al*. (2003b) that were based on estimates of a static structural model of female labor supply and program participation reveal a substantially larger effect of the 1999 in-work benefit reform, with an estimated increase in lone mothers' employment of 5 percentage points. An employment response of similar size is reported in the study by Gregg and Harkness (2003), which uses a difference-in-difference estimation technique combined with propensity score matching.

Our paper contributes to this literature by providing new evidence on the labor market effects of WFTC for lone mothers using longitudinal data drawn from the British Household Panel Survey (BHPS) during the 1991-2001 period. To isolate the diverse work incentives induced by the reform, we examine its impact both on the decision of working 16 hours per week or more and on the decision of working 30 hours per week or more, and we also estimate the impact on monthly earnings. By using longitudinal data, not only are we better able to control for changing sample composition over time, but we can also assess whether the introduction of WFTC led to changes in the rate at which lone mothers entered and left the labor force with year-to-year employment transitions. The richness of the BHPS data allows us to consider a variety of additional outcomes — such as FC/WFTC receipt, childcare usage and

[7] The labor supply response of men and women in couples is instead ambiguous. While the incentives for men (or primary earners) are to move into work, those of secondary earners are to move out of work altogether (Blundell *et al*., 2000). The effects on other aspects of household behavior (e.g., fertility, marriage stability, and childcare usage) are also ambiguous. While an analysis of WFTC's impact on couples is equally important, in this paper we limit our attention to lone parents.

expenditures, entry into marriage and further fertility — many of which have never been analyzed before in the British context. Examination of such responses however is important both because it provides us with 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 (e.g., higher incentives to form single-mother households and lower incentives to enter into marital unions), which may be crucial for the longer-term success of the reform itself.

We identify the WFTC effects on single mothers' behavior through the differential tax and benefit treatment that they receive as compared to single women without children. Our eleven-year panel regression approach improves on most of the past research that compares outcomes of different groups of women drawn from cross-sectional datasets, collected at two or at most three points in time. In fact, in our methodological analysis and in estimation, we emphasize the importance of controlling for demographic differences between women in those two groups as well as changes in characteristics, which can happen at different points in time (both before and after the reform took place) and at different stages over women's life cycle.

Our estimates suggest that financial incentives have powerful effects on lone mothers' employment and other decisions. The introduction of WFTC is estimated to have led to an average increase of about 7 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 (30 or more hours per week). With approximately 1.5 million lone-mother households, our finding means that an additional 135,000 lone mothers were in eligible employment by the end of 2001 as a result of the WFTC reform. 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. With an estimated average increase in gross earnings of about £67 per month among workers, those figures amounted to a net increase in the total wage bill of the order of £130 million between the introduction of the reform and the end of 2001. The average employment effect however conceals considerable variation in responses, which varied between 12 percentage points for lone mothers with one pre-school aged child to essentially no effects for mothers of multiple older children. We also find evidence of important (and perhaps unintended) effects on lone mothers' behaviors other than on employment. In particular, the reform led to a significant reduction in single mothers' subsequent fertility and in the rate at which they married.

This paper improves on past research in two additional respects. First, we try to identify which policy parameters accounted for the large and heterogeneous estimated employment responses and find a great deal of evidence pointing to the role played by the generous childcare tax credit component of WFTC. More than 50 percent of the increased entry rate in eligible employment was attributable to lone mothers who also chose paid childcare arrangements, and the effect was stronger for mothers with pre-school aged children. Similarly, among single mothers who continued to be in employment, 50 percent of their greater post-reform labor market attachment is accompanied by the use of paid childcare services. Second, unlike most previous evaluation research, our analysis allows explicitly for behavioral responses in the period following the announcement of the reform up to its actual implementation. We find the strongest behavioral responses in anticipation of the WFTC reform for lone mothers who were expected to benefit the most, and who indeed showed the strongest responses following the implementation of the WFTC reform. We find that the timing of such responses corresponds closely to those of the government's announcements and press/media coverage of the planned introduction of WFTC, and we are able to rule out other policy changes as potential causes for this observed behavior. Our estimates indicate that while the employment increase in anticipation of the reform was accompanied by a temporary increase in unpaid childcare arrangements, the post-reform increase in employment was instead accompanied by a large increase in paid childcare use. Taken together, these results suggest that lone mothers adjusted their behavior in anticipation of the WFTC reform, and point to the importance of modelling such effects in future evaluation research.

The remainder of the paper is organized as follows. Section I explains the eligibility rules and structure of the FC/WFTC programs, and places them in the context of other relevant labor market and welfare reforms that occurred around 1999. Section II introduces the data and describes the outcomes and the other variables used in our analysis. Section III outlines some key methodological issues related to non-experimental policy evaluations and discusses our identification strategy and estimation approach. Section IV presents the empirical results for one specific labor market outcome, that is, working 16 hours per week or more. It also distinguishes the effects by child's age, examines employment transitions, and links our estimates to those in the literature. Section V considers the results for the other outcomes, while Section VI investigates whether the responses observed in 1998, the year prior to the introduction of the reform, can be reliably interpreted as anticipation effects. Section VII explores some potential explanations for the large employment responses to the WFTC reform, especially among mothers of young children. We focus on changes in two WFTC parameters, the child credit component and the childcare tax credit. Section VIII summarizes our main results.

I. The WFTC Program in the Context of Other Recent Welfare Reforms in the United Kingdom

A. The In-Work Benefit Reform

Up to April 2003, the main in-work support program in the UK has been the Working Families' Tax Credit (WFTC), which replaced Family Credit (FC) on October 5th, 1999.[8] Along with other active labor market programs, such as the various welfare-to-work "New Deal" schemes,[9] 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. 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

[8] In April 2003, WFTC was itself replaced by two new tax credits: the Child Tax Credit (CTC) and the Working Tax Credit (WTC). CTC is an income-based credit for low-income and middle-income families who are in or out of work and who have responsibilities for at least one child under the age of 16 (or under the age of 19 if in full-time education). WTC is an income-based credit for working adults, whose eligibility criteria for families with children are similar to those of WFTC (see below and Table 1). We are not in this paper able to estimate the effects of CTC and WTC, since their introduction has occurred after the end of our sample period.

[9] The current New Deal schemes (for young people, for people aged 25 and over, for people aged 50 plus, for partners, for disabled people, and for lone parents) are directed towards individuals in different age groups and in different marital conditions, but in similar, relatively poor, economic circumstances. Although different, each scheme aims at helping people find work or return to work by offering assistance with job search, training and childcare. See also below for further details on the New Deal for Lone Parents.

family) must work 16 hours or more per week. At its introduction in 1988, Family Credit set the minimum hour cutoff at 24 hours per week, which was reduced to 16 in 1992.[10] 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.

The amount of the weekly credit to which a family is entitled depends on the number and ages of children, net family income, hours worked, and childcare costs. These affect both earned income and maximum credit. For example, in 1999 under the WFTC regime, 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, 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 constant 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 more than she would in 1998 (£53.65 versus £21.25 per week). These figures can be computed using the parameters listed in Table 1, which summarizes the main parameters of FC/WFTC over the history of the program that overlaps with our sample period.

There are five 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

[10] 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).

between 1991 and 1998 was, on average, 6.6 percent only).[11] 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, any payment of child maintenance received by a divorced mother (living alone or with another partner) was fully disregarded under WFTC, whereas under FC the weekly maintenance disregard was only £15 per child. Fourth, the taper rate (tax rate on earnings) by which the maximum credit is reduced was lowered from 70 percent to 55 percent. Fifth, WFTC is more generous with eligible childcare costs.[12] From October 1994, FC allowed eligible childcare costs (up to a maximum amount, which was £60 per week just before WFTC was introduced)[13] to be disregarded from the calculation of net family income.[14] Under WFTC instead, these costs were added to the maximum credit, and were 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.

Unlike FC, WFTC was not administered by the Benefits Agency but by Inland Revenue.[15] 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).[16]

[11] FC was generally more favourable to older children. For example, the weekly rates for children aged 11-15 in 1997 was £7.90 greater than the rates for younger children (aged 0-10). This differential treatment was reduced in November 1998 to £5.60 per week. But under WFTC the variation by age was progressively eliminated. In October 1999, the child credit for children aged 0-10 rose to £19.85 per week, only £1.05 less than the credit for children aged 11-15. In 2001, the difference was even smaller at £0.75.

[12] 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.

[13] This was the disregard for families with one child. In 1998 a disregard of £100 was introduced for families with two or more children.

[14] Gregg and Harkness (2003) report that the childcare disregard under FC was never widely used.

[15] Similarly to FC, the size of the WFTC award was assessed on weekly earnings for new claimants, while for claimants with stable jobs it was calculated by looking at the past four pay checks (or seven if paid weekly). The WFTC amount then was paid biweekly or monthly at the same rate for six months regardless of any change in income. Importantly, these features distinguish FC/WFTC from EITC, which operates as an annual tax rebate (Brewer, 2001; Blundell and Hoynes, forthcoming).

[16] 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.

It is important to note that this reform did not come unexpectedly. Since the May 1997 general elections and the Budget of July 1997, the (Labour) government was openly committed to a strategy of tax and benefit reform, in which a new in-work benefit was anticipated to be introduced drawing directly upon the experience in the United States of the Earned Income Tax Credit (Strickland, 1998). With the Pre-Budget Statement in November 1997, the government announced that a new tax credit for working families would be one fundamental element of its welfare-to-work strategy, along with a minimum wage and a general reform of welfare benefits. Although a few aspects of the reform were not publicized and perhaps not even known at that point in time, the 1998 Budget speech (delivered in March) set out the main features of the new Working Families' Tax Credit, which was to replace Family Credit in October 1999. In our analysis we therefore need to consider the possibility that employers and workers may have adjusted their labor market and other behaviors well before the actual introduction of WFTC. For example, in a slack labor market workers may decide to remain in their jobs rather than quit in order to be able to benefit from the anticipated increase in in-work support. The expectation of higher future in-work benefits could also lead those searching for a job to increase their search intensity and increase their job acceptance rates. It is precisely for these reasons that in the empirical analysis below we will allow for the possibility of behavioral responses that predate the actual implementation of the reform. We shall refer to such responses as 'anticipation effects'.

B. Other Programs, New and Old

The introduction of WFTC was also accompanied, preceded and followed by the introduction of a number of new programs and by changes in key parameters of other existing schemes. Among the new programs, perhaps the two most relevant ones are the National Minimum Wage (NMW) and the various New Deal schemes. In April 1999, a NMW was introduced as a way of reversing the rising trend in wage inequality over the previous 20 years (Dickens and Manning, 2002; Stewart, 2004).[17]

[17] Initially the rate was set at £3.60 per hour for individuals aged 22 or over and at £3.00 for those aged 18-21 inclusive. No minimum was specified for individuals aged less than 18. The adult rate was raised to £3.70 in October 2000, to £4.10 in October 2001, and to £4.20 in October 2002. The development rate (for people aged 18-21) increased to £3.20 in June 2000, to £3.50 in October 2001, and to £3.60 in October 2002. Prior to 1999 the UK had no minimum wage except for Wage Councils, which set minimum wages for a number of low-pay sectors before they were abolished in 1993. A minimum

Interestingly, in many policy discussions that prepared the reform of Family Credit, WFTC was motivated at least in part by the possibility that it could reinforce the effect of the NMW by making work pay (H.M. Treasury, 1997).

The other considerable change was induced by the New Deal (Blundell, 2002; Blundell *et al.*, 2002; Van Reenen, forthcoming). Between July 1997 and October 1998, the government launched New Deal programs aimed at six different groups of people. These comprised: a) young people (aged 18-24) who had been unemployed and received Jobseeker's Allowance (JSA) for six months or more; b) people aged 25 or more who had been unemployed and claimed JSA for 18 months; c) people in couples (aged 25 or more without children, or aged 18-24 with children) whose partner had been on JSA for six months; d) people aged 50 and over who had received a qualifying benefit (e.g., Income Support, Incapacity Benefit, and JSA) for six months or more; e) disabled people who received a qualifying benefit (such as Incapacity Benefit, Severe Disablement Allowance, or Income Support including a disability premium); and f) all lone parents in receipt of Income Support with children under 16 and whose youngest child was over five years and three months (from April 2000 this lower age cut-off was dropped to three).[18] Although different, all six schemes were intended to help low-income people move from welfare into work using a combination of intensive job-search assistance and small basic skills courses.[19] For example in the New Deal for Lone Parents (NDLP), lone parents were assigned to a personal advisor, whom they were supposed to meet once every two weeks to receive advise on job vacancies, in-work benefits, childcare arrangements, training, and job search techniques.[20] One interesting aspect of NDLP, which is largely shared with the other programs except for the New Deal for Young People and the New Deal for People aged 25 and over, is that involvement in the scheme and

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wage also existed in the agricultural sector, but that accounted for about 1 percent of total employment (Dickens and Manning, 2002).

[18] Eligibility to and provisions of the various New Deal schemes have slightly changed over time. In relation to the New Deal for Lone Parents (NDLP), since 2002 lone parents are eligible to NDLP not only if they are in receipt of Income Support (as they were in previous years) but also if they receive other benefits (including Incapacity Benefit, JSA, 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 to WFTC) who are not claiming any benefits, except child benefit. These features reinforce the notion that WFTC interacted with NDLP.

[19] Blundell *et al.* (2002) and Van Reenen (forthcoming) examine the labor market impact of the New Deal for young unemployed, and thoroughly describe the entire New Deal program.

[20] For lone parents there was also additional ad hoc financial help with travel costs (up to £20 a time), with childcare costs in the case they started working less than 16 hours per week (otherwise they would qualify for WFTC), and with training activities (up to £15 per week).

searching for work is entirely voluntary, and benefit entitlements do not depend on whether people decide to enter the scheme or not.[21]

A final feature of the post-1997 reforms were changes in a number of parameters of already existing social security benefits and programs. We highlight four specific changes. First, the generosity of Income Support (IS) payments to workless couples and nonearning lone parents with children under age 11 grew broadly in line with the child credit component of WFTC (see Table 1). Between 1998 and 1999, the IS allowance for children aged 0-10 increased by £5.10 per week (from £19.80 to £24.90), while the WFTC credit for children in the same age group grew by £5.00 per week. As in the case of WFTC, this increase eliminated a large fraction of the benefit variation by age of children. The almost equivalent increase for lone parents working less than 16 hours eliminated the work incentives that the increase in the child credit component of WFTC could have created. WFTC's work incentives must therefore stem from either higher thresholds, the lower tapering rate, more generous childcare credits, or a combination of these three parameters.

Second, the amount of Child Benefit (CB), which is paid to all people who are responsible for a child, remained fairly stable over the sample period, with single parents generally receiving a higher rate on the eldest eligible child than parents in couples (Table 1). However, the lone-parent rate of CB — which replaced the One Parent Benefit in April 1997 — was abolished from July 1998 (this is the reason why the figures from 1998 onward are in parentheses in Table 1). Although single parents could still claim the higher lone-parent rate if they were already receiving it before July 1998, its repeal may have increased the work incentives for new cohorts of single parents (through an income effect). In Section VI we explore the possibility that some of the employment effects we attribute to the announcement and actual implementation of the WFTC program, were in fact a result of the abrogation of the lone-parent rate of CB.

Third, the old rule that required employees to pay an 'entry fee' into National Insurance (NI) (this was 2 percent of their earnings up to the lower earnings limit) was abolished with effect from April 1999. In addition, from April 2000, employees

[21] However, Work Focused Interviews (WFI) for lone parents claiming Income Support were introduced in April 2001. WFI were compulsory, whereby people of working age seeking to claim Income Support were obliged to participate in a work-focused interview with an advisor at the start of their claim as a condition of receiving the benefit. For an evaluation of the effect of WFI on the probability of leaving social security benefits for lone parents and disabled, see Kirby and Riley (2003).

earning below a specified 'primary threshold' no longer paid NI contributions, and this threshold was aligned to the Pay As You Earn (PAYE) threshold.[22] The gains associated with these changes, and with the introduction of a 10-pence income tax band in April 1999,[23] were likely to be attractive especially for low-income workers and people (or families) working 16 or more hours per week.

A fourth aspect of the change in the benefit-tax structure was the growing importance of housing costs for low-income families who could receive WFTC (Blundell and Hoynes, forthcoming; Gregg and Harkness, 2003). Table 1 shows that the rules of Housing Benefit (HB) were left relatively unchanged and its main parameters were altered fairly smoothly over time. However, the eligible amounts — which, among other things, depend on rents and family needs or 'applicable amount' (see Table 1) — and receipts increased sharply over the period. For example, in the UK between 1988 and 1998, the average local authority rents more than doubled and the average private rents more than tripled (in the south-east of England, which includes London, the increase was even steeper).[24] These trends continued after 1999. Because individuals who worked 16 or more hours per week (and thus were not

[22] NI is one of the two compulsory deductions on most people's pay slips in the UK (the other being Income Tax). NI is deducted at source from employees' wages or paid directly to the Inland Revenue by the self-employed and other nonemployed groups. NI is used to fund the National Health Service and state benefits for people unable to work or on very low incomes. Entitlement to those benefits (and, in some cases, the amount paid) depends on the contribution record of the claimant. Each year a lower earnings limit and an upper earnings limit are set: in 1999 these were £66 and £500 per week, respectively (in 1998 they were £64 and £485, while in 2000 they were £67 and £535, and the primary threshold was set at £76). In 1998, a lone mother earning less than £64 per week would have paid no NI contribution (and thus, she would have not earned entitlement to social security benefits such as JSA, Incapacity Benefit, and state retirement pension). If she earned anything in excess of £64 per week (and less than £485), she would have paid 2 percent on the first £64 and 10 percent on the rest (up to £485). In 1999, the same woman would have again paid no contribution if she earned less than £66 per week. But if she earned more, she would have paid 10 percent on the amount exceeding £66 (and up to £500), and she would have not paid anything on the first £66. In 2000, with the primary threshold at £76, she would have paid NI contributions only if her earnings were greater than £76 per week. If her earnings were below £67 (the lower earnings limit) she would have paid no contribution (and not been entitled to benefits). So if her earnings were between £67 and £76, she would have been treated as if she had paid NI contributions on her earnings.

[23] For the first time in 1999, the government introduced a starting tax rate of 10 percent on annual incomes below £1,500. There was then a basic tax rate of 23 percent on the next £26,500, and a higher rate of 40 percent on incomes over £28,000. In 1998, instead, there were only two rates, a lower rate of 20 percent applicable on the first £4,300 and a higher rate of 40 percent on incomes over £27,100. In April 2000, the basic rate was reduced from 23 to 22 percent, while the income bands have increased to £1,520 (for the starting rate), to the next £26,880 (for the new basic rate) and to over £28,400 for the higher rate. Since then, only the income bands have changed but not the rates. See Inland Revenue (various years).

[24] The source of this information is the Office of the Deputy Prime Minister web page (under Housing Statistics and Live set of tables, Sections 6 and 7 on Social Housing and Rents and Tenancies, respectively) ⟨http://www.housing.odpm.gov.uk/statistics/live⟩.

entitled to receive IS) could not receive maximum HB if their income was greater than their specific applicable amount, their incentive to work while receiving WFTC could be offset by more generous IS and HB payments. For those in the rented sector, therefore, the potential loss of HB could counteract the work incentives associated with the 1999 in-work benefit reform.

In sum, there are a number of possible interactions between WFTC and other policy initiatives (e.g., minimum wages, New Deal programs, national insurance contributions, Child and Housing Benefits and Income Support). 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 evaluation methodology is deferred until Section III.

II. Data

The data we use are from the first eleven waves of the British Household Panel Survey (BHPS) collected over the period 1991-2001. Since Fall 1991 the BHPS has annually interviewed a representative sample of about 5,500 households covering more than 10,000 individuals. All adults and children in the first wave are designated as original sample members. On-going representativeness of the non-immigrant population has been maintained by using a 'following rule' typical of household panel surveys: at the second and subsequent waves, all original sample members are followed (even if they moved house or if their households split up), and there are interviews, at approximately one-year intervals, with all adult members of all households containing either an original sample member, or an individual born to an original sample member whether or not they were members of the original sample. The sample therefore remains broadly representative of the population of Britain as it changes over time.[25]

Our estimation sample includes unmarried non-cohabiting females (separated, divorced, widowed and never married) 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

[25] Of the individuals interviewed in 1991, 88 percent were re-interviewed in wave 2 (1992). The wave-on-wave response rates from the third wave onwards have been consistently above 95 percent. See Taylor (2003) for a full description of the dataset. Detailed information on the BHPS can also be obtained at ⟨http://www.iser.essex.ac.uk/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.

ill or disabled, or in school full time in a given year.[26] The sample includes 3,333 women who at some point during the observation period where living alone, of whom 1,507 are lone mothers and the remaining 1,826 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 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 October, 5th. To limit problems of interpretation, they were dropped from the sample analyzed in this paper. Their inclusion however does not alter any of our main results.

Table A1 presents summary statistics of the labor market outcomes and characteristics of the two groups of women. The first column presents the statistics for all unmarried women without children, while the second column presents the statistics for all unmarried women with children. There are some noticeable differences in characteristics between the two groups. Those who have children tend on average to be younger (28.5 versus 31.1 years),[27] less educated (64 percent have O-level or lower qualifications versus 46 percent among childless women, while only 4.5 percent of lone mothers have a university degree versus 14.3 percent), more likely to be nonwhite (8.2 versus 4.3 percent), and more likely to be in social housing (35 versus 20 percent). In addition there appear to be systematic differences in the employment behaviors of both groups of women. Compared to unmarried childless women, lone mothers are less likely to work 16 or more hours per week (41 versus 64 percent), and have a lower probability of staying in such labor market state (0.65 versus 0.91 probability) and entering it (0.20 versus 0.27 probability) in any given two successive years.[28] They are also less likely to work 30 or more hours per week (26 versus 53

[26] Eissa and Liebman (1996) use similar sample selection criteria.
[27] Because single childless women tend to be more concentrated at the bottom and top ends of the age distribution, there may be important age effects. In our empirical analysis we account for such effects parametrically by using a quartic polynomial in age.
[28] Throughout the paper, worked hours are defined by usual weekly hours of work plus usual weekly hours of overtime work.

percent), work any positive number of hours (60 versus 73) and, even after conditioning on working positive hours, their average monthly earnings are approximately 40 percent lower.[29] Finally, they are more likely to be in receipt of Income Support (39 versus 8 percent). The other outcomes listed in Table A1 are relevant only for single mothers (FC/WFTC receipt and amount awarded, usage of and expenditures on childcare services, entry into marriage, and having an additional child), while entry into lone motherhood is an outcome that we analyze for single childless women only.[30]

Figures 1-3 plot the time trends for all outcomes between 1991 and 2001 distinguishing again between single mothers and single women without children, where it is appropriate. Figure 1 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 but only up to 1998, when they began to rise from about 40 to nearly 48 percent.[31] Figure 1, Panel (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. To analyze whether this increase was due primarily

[29] Conditioning at greater levels of labor supply leads to lower earnings differentials, although these are still sizeable. If we condition on working 16 or more hours per week, lone mothers earned on average £820 per month (2001 prices) while single women without children earned £1038 (about 27 percent more). If we condition on working 30 or more hours, lone mothers earned £1,030, approximately 24 percent less than their single childless counterparts.

[30] Most of the figures in Table A1 conform to official statistics and to those reported in related studies (e.g., Blundell *et al*., 2000; Gregg and Harkness, 2003). Perhaps the two most notable exceptions are childcare (usage and expenditures) and FC/WFTC award. 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 A1), 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 A1 are computed over families where the youngest child is aged 12 or less. (In fact, the BHPS collects childcare information only from households that have at least one dependent child aged 12 or less). Official statistics show that the average FC/WFTC award for lone parents over our sample period was £64.82 per week (Inland Revenue, 2003), implying that the BHPS data underestimate average FC/WFTC awards by approximately 35 percent.

[31] 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 unusual changes.

to an increase in the rate at which women entered eligible employment or to a reduction in the rate at which they left it, we plot both rates in Figure 2. Panel (a) shows a relatively stable persistence rate until 1998, after which this rate increased for lone mothers from approximately 60 percent to 70 percent. Similarly, in Panel (b) we see an increase in the rate at which single mothers entered eligible employment, beginning in 1998, especially when compared to the rate for single women without children.

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 1(a). The evolution of earned income for lone mothers shows a slope change in 1999, whereas the trend for unmarried females without children has been relatively stable since 1994, with the possible exception of 1996 (Figure 3, Panel (b)). Although the real FC/WFTC awards to lone mothers continued to rise from 1993 to 1999 and thereafter (Figure 3, Panel (d)), the proportion of lone mothers in our sample receiving the credit remained fairly constant between 1991 and 1998, but substantially increased from 45 percent in 1998 to almost 60 percent in 1999 (Figure 3, Panel (c)) and kept on growing after 1999, albeit at a slower pace. Conversely, IS receipt continued to fall in our sample of lone mothers, and somewhat more rapidly in conjunction with the 1999 reform, so that by the end of the period it was about 30 percent, almost 20 points lower than in 1992. We observe a much lower proportion of single childless women in receipt of IS, which also gradually declined over the period (Figure 3, Panel (e)). Usage of and expenditures on formal childcare services by lone mothers were stable up to 1998, and increased only in concomitance with the WFTC reform (Figure 3, Panels (f) and (g)). Finally, the 1999 reform appears to be associated with a reduction in the entry rates into marriage and new additional births for lone mothers (Figure 3, Panel (h)) as well as with a small decline in the entry rate into lone motherhood (Figure 3, Panel (i)).

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. They, and the employment transition rates in particular, also suggest that there were behavioral responses in anticipation of such a reform. To investigate the causal link between the introduction of WFTC and these socioeconomic 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. The next section discusses the methodology we use to identify the responses of British lone mothers to the WFTC reform.

III. Evaluation Methodology

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 Angrist and Krueger [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 the United Kingdom 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 in-work benefit reform occurs (i.e., s=1999). Suppose the outcome variable y_{it} (e.g., labor market participation or full-time employment) is determined by the following specification

(1) $$y_{it} = \alpha_1 + \alpha_2 d_i + \alpha_3 t + \alpha_4 \mathrm{I}(t \geq s) + \beta d_i \mathrm{I}(t \geq s) + u_{it}$$

with $\mathrm{E}(u_{it} \mid d, t)=0$ where $\mathrm{E}(\cdot)$ is the mathematical expectation operator.

In equation (1), which for the moment excludes individual characteristics, α_3 reflects a linear time trend common to both lone mothers and single women without children, and the term $\mathrm{I}(w)$ is a function indicating that the event w occurs, so that β is the parameter that captures the treatment effect (i.e., the WFTC effect). The parameter α_4 represents a shift in the average value of y_{it} that is shared by both groups of women. In our case, this captures the effect of all the other (non-WFTC) policy changes that occurred at s as discussed in Section I (e.g., the introduction of the minimum wage and the income tax rate change). The residual u_{it} is assumed to be an i.i.d. term.

Note that when $\alpha_4 = 0$ and the sample contains data on at least three different periods, a control group is not essential in this regression approach. So without a structural change also for the control group, the treatment effect could be identified by simply exploiting the time variation in the outcome for lone mothers before and/or after the introduction of WFTC in combination with a linear trend assumption. Allowing for such a structural change, however, is likely to be important in our case because, as we discussed in Section I, the introduction of WFTC was part of a larger set of reforms of the tax-benefit program that took place in 1999. 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 reforms that took place in that year. By assuming that lone parents would have responded in the same way to these reforms, we are able to net out the separate impact of WFTC.

A widely used approach to identify β is the "difference-in-difference" (DD) method (Card, 1990; Eissa and Liebman, 1996; Angrist and Krueger, 1999; Meyer and Rosenbaum, 2001; Bertrand *et al.*, 2004). The DD estimator is usually based on data from two periods

$$(2) \qquad \mathrm{DD} \equiv \left[\mathrm{E}(y_{i,s-1+k} \mid d_i = 1) - \mathrm{E}(y_{i,s-1-k'} \mid d_i = 1)\right] \\ - \left[\mathrm{E}(y_{i,s-1+k} \mid d_i = 0) - \mathrm{E}(y_{i,s-1-k'} \mid d_i = 0)\right],$$

where $s-1+k$ and $s-1-k'$ (with $k>0$ and $k' \geq 0$) respectively represent the post- and pre-reform periods. In the case of model (1) this method identifies the treatment effect β.[32] As it is based on data from only two periods, this approach requires a control group even when $\alpha_4 = 0$.

A well-known limitation of this specification is that it only allows for a common trend captured by α_3 (Meyer, 1995). But if, say, the employment rates of lone mothers evolved differently from those of single women without children regardless of the in-work benefit reform, then a better specification for y_{it} will be given by

$$(3) \qquad y_{it} = \alpha_1 + \alpha_2 d_i + (\alpha_{31} + \alpha_{32} d_i)t + \alpha_4 \mathrm{I}(t \geq s) + \beta d_i \mathrm{I}(t \geq s) + u_{it}.$$

In this case, the DD method yields a treatment effect estimate equal to $\beta + \alpha_{32}(k+k')$, where $k+k'$ represents the average number of calendar periods (years or months) between the post-reform and pre-reform period observations in the sample. Unless $\alpha_{32} = 0$ this is clearly a biased estimate of β. The bias arises precisely

[32] Although DD and regression-based estimators when applied to the same data will provide the same point estimate, their standard errors will generally be different, because the former allows the error variance to be different for the pre- and post periods and for each group.

because the time evolution of the outcome variable differs between control and treatment groups. In fact, Figures 1-3 show that single childless women exhibit a number of labor market patterns that are different from those of lone mothers. However, under specification (3) it is easy to show that a "difference-in-difference-in-difference" (DDD) estimator, defined as

$$(4) \qquad \text{DDD} \equiv \{[\text{E}(y_{i,s-1+k} \mid d_i = 1) - \text{E}(y_{i,s-1} \mid d_i = 1)] - [\text{E}(y_{i,s-1} \mid d_i = 1) - \text{E}(y_{i,s-1-k} \mid d_t = 1)]\}$$
$$- \{[\text{E}(y_{i,s-1+k} \mid d_i = 0) - \text{E}(y_{i,s-1} \mid d_i = 0)] - [\text{E}(y_{i,s-1} \mid d_i = 0) - \text{E}(y_{i,s-1-k} \mid d_t = 0)]\}.$$

will identify β (the time subscript k indicates the length of the time periods over which the differences are computed). Thus, in presence of different time trends in outcomes for lone mothers and single childless women, the availability of multiple observations before and/or after the occurrence of the policy change allows us to uncover the treatment effect using a DDD approach. An important, and often overlooked, point is that the estimator defined in (4) requires equal time intervals in the computation of the differences. In fact, if the lengths of the time periods differ in the pre- and post-difference terms in (4), and say the former is k' while the latter is k, then the DDD estimator will not yield β, but rather $\beta + \alpha_{32}(k - k')$, which confounds the treatment effect with the difference in trends between the two groups. The bias depends on the magnitude of $(k - k')$, which could be either negative or positive.[33]

A specification for y_{it} that is slightly more general than (3) is

$$(5) \qquad y_{it} = \alpha_1 + \alpha_2 d_i + (\alpha_{31} + \alpha_{32} d_i)t + [\alpha_{41} + \alpha_{42}(t - s)]\text{I}(t \geq s) + \beta d_i \text{I}(t \geq s) + u_{it},$$

which, in addition to different group-specific time trends and a common jump in y_{it} at s, allows for a common change in the trend at the time of the policy intervention. This means that equation (5) allows for non-WFTC shocks that might have occurred in 1999 to have a common independent effect on both intercepts and slopes of the process generating y for both lone mothers and single women without children. With this specification neither DD nor DDD estimators can identify the impact of the in-

[33] However, a simple modification of the DDD estimator in which the differences $[\text{E}(y_{i,s-1} \mid d_t)$ - $\text{E}(y_{i,s-1-k'} \mid d_t)]$ are multiplied by k/k' would yield β. The interpretation of the DDD estimates reported in Gregg and Harkness (2003) is problematic because their pre- and post-reform time intervals are of different lengths, and their estimates are not reweighted.

work benefit reform. In fact, the DD method yields $\beta + \alpha_{32}(k + k') + \alpha_{42}k,$ whereas the DDD method yields $\beta + \alpha_{32}(k - k') + \alpha_{42}k$.[34]

Our evaluation strategy will be based on specification (5), which we extend in a number of ways. First, to control for potential differences in group-specific compositional changes over time, we include a set of standard individual characteristics (e.g., age, education, region of residence, and number and age of children). Second, because we use panel data, we also account for compositional changes in unobserved characteristics by allowing for 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 are endogenous to the policy reform.[35] These two extensions lead to the following panel-data regression model:

$$(6) \qquad y_{it} = \alpha_1 + \alpha_2 d_{it} + (\alpha_{31} + \alpha_{32} d_{it})t + [\alpha_{41} + \alpha_{42}(t-s)]\mathrm{I}(t \geq s) + \beta d_{it}\mathrm{I}(t \geq s) + \mathbf{X}'_{it}\gamma + \theta_i + \varepsilon_{it},$$

where $\mathbf{X}_{it}$ is the vector of individual characteristics, θ_i represents the individual fixed effects, and ε_{it} is an i.i.d. error term.

So far we have treated the WFTC reform as an event that took place at a specific point in time with an immediate and permanent impact on people's behavior. In Section I however we emphasized the relatively long build-up to the reform through the Treasury's Pre-Budget of November 1997 and the highly publicized Budget of March 1998. Lone mothers' behavior therefore may have been affected even before the official introduction of WFTC in October 1999. For instance, in markets where jobs are hard to find, a nonworking lone mother could search for a job more intensely or accept job offers more easily in anticipation of the greater future benefits offered by WFTC. Similarly, expectations of increased benefits may reduce the rate at which working single mothers quit their jobs. Anticipation effects of this type have been largely ignored in the evaluation literature (Abbring and Van den Berg [2003] make explicit the assumption of no anticipation effects that underlies most evaluation research; while Attanasio *et al.* [2001] in their structural analysis of the

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[34] If the length of the time periods is the same for the pre- and post-difference terms, then $k = k'$ and differences in the time evolution of the outcome variable between treatment and control groups do not contribute to the bias.

[35] Another advantage of using longitudinal data is that, by avoiding the sampling variability present in repeated cross-sectional data, we substantially increase the statistical power for testing and estimating the policy effect.

Progresa program in Mexico allow the control group to anticipate being included into the program at a future date, and find some mixed evidence). Thus we extend model (6) and test for the presence of such anticipation effects by allowing a nonzero treatment effect in 1998.

Closely related to anticipation effects is the issue of timing and duration of the program impact. There could be delayed responses if, for example, it took time for women to find eligible employment or there could be learning about the program benefits, which could be greater or smaller than individuals initially anticipated. Furthermore, the impact of WFTC could have interacted with labor market trends and other changes in the post-reform period. We incorporate these dynamic effects as follows:

$$(7) \qquad y_{it} = \alpha_1 + \alpha_2 d_{it} + (\alpha_{31} + \alpha_{32} d_{it})t + \left[\alpha_{41} + \alpha_{42}(t-s)\right] I(t \geq s) + \sum_{t=1998}^{2001} \beta_t d_{it} + \mathbf{X}'_{it}\gamma + \theta_i + \varepsilon_{it} ,$$

where β_t represents the treatment effect at time *t* and β_{1998} is the anticipation effect.

By following the same individuals over time, the BHPS data allow us to 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. By relating changes in employment transitions to employment levels, we can better gauge the presence of anticipation effects as well as assess the lasting effects of the reform over and above its immediate impact. We implement this analysis by estimating separate outcome equations (7) for each value of $y_{i,t-1}$.

Before turning to the results, it is useful 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 of the true effect.

IV. Econometric Evidence on Working 16 or More Hours per Week

A. Main Estimates

Tables 2-4 present the estimates for our main labor market outcome, that is, working 16 or more hours per week. For simplicity of interpretation of the coefficients, we present results based on ordinary least squares. Logit (and Chamberlain fixed-effects logit) estimates provide qualitatively similar results (these results are available from the authors upon request). Table 2 shows treatment effect estimates for a number of specifications of models (6) and (7). We estimate models with and without individual fixed effects θ_i (fixed-effects and level estimates, respectively). The figures in column (i) are obtained from regressions that impose a constant treatment effect $(\beta_{1999} = \beta_{2000} = \beta_{2001} = \beta)$ and exclude any possible pre-program trend $(\alpha_{31} = \alpha_{32} = 0)$. Different pre-program trends are instead included in specification (ii), while specification (iii) also incorporates year-specific treatment effects. For each specification we also consider the case that allows for anticipation effects.

Focussing on the fixed-effect estimates, we find that the rate at which lone mothers worked 16 or more hours per week increased by a statistically significant 5.3 percentage points. Allowing for different pre-program trends reduces this effect only slightly to 4.7 percentage points. This estimate is remarkably close to those reported in Brewer *et al*. (2003b) and Gregg and Harkness (2003), who applied different methods to different data from ours.[36]

Table 2 also reveals evidence of substantial anticipation effects (in 1998), of the order of 4 percentage points. Accounting for such effects leads to systematic increases in all the post-implementation estimates, which range from about 1.5 to almost 3 higher percentage points. For example, in the case of specification (ii), an anticipation effect of 4.4 percentage points is accompanied by an overall post-1998 impact of 7.3 points. This effect is approximately 50 percent greater than the effects reported in the two above-mentioned British studies, and is also at least three times larger than the effect found by Blundell *et al*. (2000) and by Blundell and Hoynes

[36] It is worthwhile noting that, in the case of specification (ii) for the basic model, the estimates of α_{32} and α_{42} are around 0.0015 (s.e.=0.0024) and 0.017 (s.e.=0.007) respectively. Therefore, following our analysis in Section III, neither DD nor DDD estimators are appropriate to retrieve the treatment effects of interest here. The point estimates of α_{32} and α_{42} and their standard errors change only slightly when anticipation effects or year-specific treatment effects are accounted for.

(forthcoming). Although the effects emerging from the level estimates are qualitatively similar to those just discussed, they are generally greater, suggesting that ignoring time-invariant unobserved effects produces a small upward bias. Specification (iii) shows that the largest WFTC employment impact emerged in 1999 (that is, immediately after its introduction), when it led to an increase in the employment rate of 9 percentage points. The employment effect fell slightly in subsequent years to almost 5 points in 2000 and nearly 7 points in 2001. Our finding of a program-introduction effect is consistent with the results reported in Blundell (2001b).[37]

B. Effects by Child's Age

By eliminating the differential treatment that Family Credit had on children of different ages (achieved via a larger credit increase in favour of younger children) and by providing more generous support to childcare costs, the WFTC reform could have generated different labor market responses depending on the number and ages of children. We explore this possibility by estimating separate treatment effects by the number of dependent children in three age intervals, $N_{it}^{(j)}$, where *j*=1,2,3 represents ages 0-4, 5-10, and 11-18 respectively. That is, in equation (7) we interact the treatment status indicator d_{it} for each of the post-1997 years with each of the $N_{it}^{(j)}$ terms and $\mathrm{I}[N_{it}^{(j)} > 0]$ dummy variables:[38]

$$(8) \qquad y_{it} = \alpha_1 + \alpha_2 d_{it} + (\alpha_{31} + \alpha_{32} d_{it})t + \left[\alpha_{41} + \alpha_{42}(t-s)\right]\mathrm{I}(t \geq s)$$
$$+ \sum_{t \geq 1998}^{2001} \left(\sum_{j=1}^{3} b_t^{(j)} \mathrm{I}[N_{it}^{(j)} > 0] + \widetilde{b}_t^{(j)} N_{it}^{(j)} \right) d_{it} + \mathbf{X}_{it}' \gamma + \theta_i + \varepsilon_{it} \,.$$

The results of our most general specification (with anticipation effects and individual fixed effects) are in Table 3, which reports estimates for the presence of one child or two children in each of the three age groups.[39]

[37] The large program-introduction effects may be related to the way WFTC was administered (see Table 1 and our discussion in Section I).

[38] Thus, in the case of one child in age group *j*, the total treatment effect obtained from (8) is given by $\beta_t^{(j)} = b_t^{(j)} + \widetilde{b}_t^{(j)}$, while in the case of *k* children in age group *j* it is given by $\beta_t^{(j)} = b_t^{(j)} + k\widetilde{b}_t^{(j)}$. Similarly, one could compute such effects for mothers with children in different age groups.

[39] When anticipation effects are not accounted for, the post 1998 estimates (not reported for brevity) show similar patterns by child's age, number of children, and post-reform year, although, as in Table 2, they are always smaller than those reported here. In some cases the reduction is substantial: for example, for mothers with one child aged 0-4, the downward bias is of the order of 44 percent.

The strongest increase in the probability of working 16 or more hours per week 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 finding larger, albeit more modest, employment effects for mothers with younger children. A lone mother with one child aged 0-4 increased that probability by 12 percentage points, and a lone mother with one child aged 5-10 experienced an increase of about 8.5 percentage points. But for a single mother with one child in the oldest group, that increase was of the order of only 4.5 percentage points. As found earlier, the reform had its largest impact in 1999 (between 7.5 and 13 percentage points, from the oldest to the youngest group respectively) and was followed by smaller effects in the two subsequent years, regardless of child's age. Lone mothers with a greater number of children generally had a much smaller labor supply response, and in most cases it is not statistically significantly different from zero. One notable exception is for mothers of two children aged 11-18, who increased their probability of working 16 or more weekly hours by approximately 6 percentage points in the entire post-reform period, and by 7 percentage points just in 1999.

Table 3 shows substantial anticipation effects, especially for mothers with children aged 0-10. For a lone mother with one child in the 0-4 age group, the anticipated effect of WFTC (in 1998) is estimated to be a 6-percentage point increase in the probability of being in eligible employment. For a lone mother with one child aged 5-10, this effect is about 5 percentage points, but for a lone mother with one child in the oldest group there is virtually no effect. Thus, the overall pattern of such anticipation effects matches well with that of the post-reform impact. Similarly, for lone mothers with more than one child, the anticipation effects are always smaller, of the order of 2-percentage point higher probability, and are not statistically significant.

A stronger labor supply response for lone mothers with very young children could be attributed to different components of the WFTC reform. There are possibly five candidates. The first is the reduction in the taper rate (i.e., the rate by which the maximum credit is reduced). This, however, affected all lone parents equally, independently of the age of their children. Two other possible sources are the child maintenance disregard and the new way in which the tax credit was administered. But again both were child-age neutral. A fourth explanation relates to the large increase in the child-specific component of the tax credit. As shown in Table 1, 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.[40] The last explanation is based on the increased generosity 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. Section VI will return to this issue in greater detail.

C. Employment Transitions

We now examine the impact of WFTC on year-to-year employment transitions. For both equations (7) and (8), we estimate the WFTC 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 latter as the entry probability. The corresponding treatment effect estimates are shown in Table 4, in which we only report estimates for the most general specification which allows for different pre-program trends. On average, the introduction of the in-work benefit reform increased lone mothers' persistence rates by 7 percentage points. Note that this increase actually began with a 4-point jump in 1998, and stabilized after 1999. Entry rates into WFTC-eligible jobs show similar patterns. They rose on average by almost 6 percentage points over the post-implementation period (with their largest increase noted immediately after the introduction of the reform), and displayed an anticipation effect of the order of 3-3.5 percentage points, albeit not statistically significant at conventional levels.

Finally, after distinguishing the three usual child's age groups (ages 0-4, 5-10, and 11-18), Table 4 shows the treatment effect estimates of persistence and entry rates for mothers with exactly one child in each of the groups.[41] In line with our previous results, the largest impact of the WFTC reform on employment persistence and entry

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

[41] We chose not to present the estimates for two children because, as suggested by Table 3, most of the statistically significant effects were concentrated in the case of mothers with one child.

rates emerged for women whose children were in the youngest group. The average persistence rate for a single mother with one child aged 0-4 increased by about 12 percentage points relative to the rate of a corresponding single childless woman. This effect declined to just less than 7 percentage points for a mother with one child aged 5-10, and dropped to zero 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 both types of transition probabilities, there is evidence of a strong program introduction effect with the largest estimates generally emerging in 1999 immediately after the WFTC onset. We also find modest to large anticipation effects, especially in the persistence rates of mothers with one child aged 5-10 and in the entry rates of mothers with one child aged 0-4.[42]

D. Summary

We find evidence of a large WFTC effect on eligible employment for lone mothers of about 7 percentage points over the entire post-reform period. Most of this increase occurred in the form of a sizable program-introduction effect, with the largest estimates emerging immediately after the introduction of the reform. This impact was also preceded by non-negligible anticipation effects: in 1998, such effects were of the order of a 4-percentage-point higher probability of working 16 or more hours per week. With approximately 1.5 million lone-mother households, our finding means that an additional 135,000 lone mothers were in eligible employment by the end of 2001 as a result of the WFTC reform. Our estimates are higher than those reported in previous UK studies (e.g., Blundell *et al*., 2000; Brewer *et al*., 2003b; Gregg and Harkness 2003). This can be for a number of reasons. Ours is the only evaluation based on panel data. Panel data allow us to better control for changing sample composition and the use of data on multiple pre- and post-reform periods permits us to better account for various aspects of individual and aggregate behavior such as

[42] The transition probability estimates line up well with the estimates shown in Tables 2 and 3. For example, given that 41.5 percent of lone mothers were in eligible employment in 1998, it is easy to verify from the estimates in column (ii) that an additional 2.9 percent were in employment in the post-reform period through the increase in persistence rates and a further 3.5 percent were in employment through the higher entry rates. These changes lead to a total effect of 6.4 percentage points, which compares well with the corresponding 7.3 percentage point figure reported in Table 2.

group-specific trends before and after the WFTC reform, non-WFTC shocks that might have occurred at the time of the reform and might have had independent effects for lone mothers and single women without children. Moreover, our evaluation of the impact of WFTC allows for anticipation effects. Unlike Blundell *et al*. (2000) and Brewer *et al*. (2003b), our analysis incorporates employment responses due to the reduction in net childcare costs, without assuming a fixed relationship between childcare expenditures and hours of work.

The strongest employment impact was on mothers with one child aged 0-4 and on mothers with one child aged 5-10. The effects for those two groups of women were of the order of 12 and 9 percentage points respectively. Although some significant effects (of the order of 6 points) emerged also for mothers living with two children aged 11-18, the labor supply effect for single mothers with two (or more) children was typically smaller. Interestingly, larger employment effects for single women with young children also emerged in the study by Gregg and Harkness (2003), and in the case of EITC, as documented by Eissa and Liebmen (1996) and Meyer and Rosenbaum (2001). Finally, the analysis of transition probabilities reveals that both greater rates of entry and greater rates of staying in jobs involving 16 or more hours per week were responsible for the large estimated effects. To identify the specific aspects of the WFTC reform that were responsible for its impact, the next section will document the impact that the reform had on a broad range of other socioeconomic outcomes for lone mothers.

V. Empirical Evidence on Other Behavioral Outcomes

We consider eleven additional outcomes. To ease our discussion, Table 5 only contains fixed-effects estimates of β and β_{1998} (column (i)) and of β_{1999}, β_{2000}, β_{2001}, and β_{1998} (column (ii)). Similarly, Table 6 shows the fixed-effects estimates of β for having one child in each of the three age groups considered above (for convenience the estimates of β_{1998} are not shown, but are similar to those reported in Table 5).

A. Full-Time Employment

As argued in Blundell *et al*. (2000) and Blundell and Hoynes (forthcoming), in absence of the childcare subsidy the increase in net income induced by the introduction of WFTC can be expected to be small below 25 hours of work per week, primarily because of the interaction between WFTC and the Housing Benefit system (see Table 1 and the discussion in Section I). At higher hours levels, however, the reduction in the WFTC taper rate is expected to have led to greater returns to working. We found earlier that the WFTC effect on the probability of working 16 or more hours was substantial, but we do not know whether this was mainly due to an increase in full-time work.

Using equations (7) and (8), we investigate this possibility by estimating fixed-effects linear probability models of working 30 or more hours per week.[43] The results from this exercise, which are in the first row of Tables 5 and 6, are remarkably similar to those reported above, suggesting that in fact the positive labor supply response of single mothers is predominantly driven by an increase in full-time employment. The rate at which lone mothers worked full time increased by almost 7 percentage points over the post-reform period. This increase was preceded in 1998 by a smaller jump of about 3.5 percentage points (which is statistically significant at the 10-percent level). There is again evidence of a program-introduction effect, with the largest increase of 9 percentage points occurring in 1999 (Table 5). All single mothers with only one dependent child experienced a steep rise in their probability of working full-time (Table 6). But mothers with one child aged 0-4 increased their chances by almost 15 percentage points, an increase that was at least twice as large as that for mothers with one child aged 5-10 or 11-18.

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 already working, Tables 5 and 6 also show estimates of the impact of the WFTC on the labor force participation rate. They indicate that a large proportion of the increase in employment was due to the entry of single mothers into the labor market. This was particularly the case for 1998 which saw an increase in employment of almost 5 percentage points. Combined, the estimates indicate that while the anticipation effect was mainly due to single mothers entering the labor force, the post-reform effects included a large proportion of women

[43] As Table 1 shows, since 1995 FC provided an additional credit for working 30 hours per week or more. WFTC continued offering this extra credit, although its generosity did not substantially change, having increased only by 35 pence between 1998 and 1999.

who were already working, who increased their hours of work in order to become eligible for WFTC benefits.

B. Labor Income

Another indication of how the reform affected lone mothers' welfare is given by its effect on monthly earnings. Conditioned on positive earnings, the least squares estimates in the second row of Tables 5 and 6 are not selectivity corrected. Among those with positive earnings, single mothers' average earnings increased by almost £67 per month (2001 prices) over the 1999-2001 period compared to that of single childless women. Therefore, the reform did not only lead to an increase in the number of working single mothers, but also led to an increase in their average monthly earnings. The estimated increase is comparable to those reported by Gregg and Harkness (2003) and Blundell *et al.* (2000). The earnings gain was evenly spread across the post-reform years with just a slight upward trend, going from £62 per month in 1999 to £70 per month in 2001. Lone mothers with one pre-school child (aged 0-4) experienced the largest growth in monthly earnings of about £126, but the wage increases for mothers with one child aged 5-10 or aged 11-18 were also substantial, and of the order of £97 and £75 per month respectively.

Interestingly, there is no evidence of an anticipation effect for this outcome. However this corresponds well to the labor supply effects discussed earlier, which suggested that the increases in eligible and full-time employment in 1998 were mainly due to more women entering or remaining in the labor force. The estimates therefore indicate that the women responsible for the increase in the labor force participation rate earned similar amounts to women already working. The large estimated post-reform earnings gains, on the other hand, reflect the increase in working hours among those working to qualify for (higher) WFTC benefits.[44]

[44] While not reported here to save space, using hours worked as dependent variable provides estimates consistent with our interpretation based on the discrete employment outcomes. More specifically, when women with zero hours of work are included, we find that the introduction of WFTC led to an average post-reform increase of about two additional hours worked per week, with a statistically significant increase of about one hour in 1998.When we consider hours worked by workers only (that is, excluding those with zero hours), all these results are confirmed, except that the 1998 effect is smaller (about 0.5 additional hours worked) and statistically not significant.

C. WFTC Receipt and Awards

High receipt rates provide a good indication of how well any public transfer program reaches its potentially eligible population. In relation to the WFTC reform, two of the government's key goals were to reduce public expenditures for achieving a given amount of poverty reduction and to reduce the stigma generally associated with income-related benefits (Section I). Therefore, even at the time of the discussions that preceded the introduction of WFTC, the policy relevance of increasing participation into the new program was apparent (Brewer *et al.*, 2003b). Because the BHPS does not collect information on individuals' WFTC entitlement, we cannot estimate take up rates. We do, however, observe FC/WFTC receipt and can therefore estimate the probability of tax credit receipt for all lone mothers in our sample. For the subsample of those receiving the tax credit, we also estimate the (selectivity uncorrected) effect of the reform on the average monetary amount received.

During the post-reform period, the probability that a lone mother received a FC/WFTC award rose significantly by 11 percentage points on average. This estimate is much greater than the change in participation rates reported in the study by Brewer *et al*. (2003b), which suggests an increase of about 3 percentage points, but it is closer to the estimated 9-point increase in take-up rates between 1998 and 2001 that emerge from published official statistics (Department for Work and Pensions, 2001; Inland Revenue, 2002). Conditional on receiving the tax credit, the average award to lone mothers went up by almost £46 per month (2001 prices). For both outcomes, there is no evidence of anticipation effects with the timing of the increase in benefits corresponding to the date benefits and eligibility actually increased. Nor is there any evidence of program-introduction effects, with the increased participation rate and award amount continuing to slowly rise after its initial 10 percentage points and £40 increase respectively in 1999. The results by child's age (Table 6) line up well with those found earlier. Again, it is single mothers with one child in the youngest age group who experienced the largest effects. Their probability of FC/WFTC receipt increased by about 17.5 percentage points (as compared to the 10-point increase for mothers with one child aged 11-18), and their monthly awards rose by £54 (while the mean monthly award grew by about £41 for mothers with one child in the oldest age group).

The lack of anticipation effects indicates that a substantial fraction of the single mothers who were responsible for the increase in eligible employment rates in

1998 (either by entering or not quitting) were either not eligible under the old FC program or did not take it up because of relatively low benefits, but became eligible after the 1999 reform. This finding points to the importance of the expansion in eligibility for receiving benefits due to the lower taper rate, the higher disregard on child maintenance payments and the more generous childcare tax credits.

D. Income Support Receipt

Another objective of the WFTC reform was to keep government costs low (Section I; Blundell, 2002). We consider one specific aspect of WFTC cost effectiveness by exploring whether or not the reform accelerated the decline in Income Support receipt observed in Figure 3(e) for both lone mothers and single women without children.

For this purpose, we estimate equations (7) and (8) using fixed-effects linear probability models of IS receipt, the results of which are in the fifth row of Tables 5 and 6. The introduction of WFTC accounted for about 25 percent of the 10-point reduction in IS participation for lone mothers, from 40 to 30 percent between 1997 and 1999. This implies that up to 37 percent of the WFTC induced employment increase, and a substantially higher fraction of the new entrants into eligible employment came from single mothers previously on IS. Interestingly, there is evidence of a strong anticipation effect in 1998, of the order of 4 percentage points. In combination with a positive and significant employment increase in 1998, these results indicate that while the initial employment increase in anticipation of the WFTC reform came primarily from single mothers on IS, subsequent employment increases came from other groups of single mothers, which is consistent with our discussion in sections A and B.

So despite the large increase in IS child allowances between 1998 and 1999 — which essentially offset the work incentives generated by WFTC through the increase in its child credit component — we observed a further reduction of lone mothers' IS participation. This finding once again emphasizes the importance of changes in other WFTC parameters (such as taper rates, applicable amounts, and childcare tax credits), as drivers of the estimated employment responses.

E. Childcare Use and Expenditures

WFTC provided much greater support for childcare than Family Credit did, in a number of ways (see section I, and Table 1). It added the childcare element towards the overall award, while under FC such costs were disregarded from income, with the result that the maximum award could not increase. Furthermore, the maximum level of support for childcare costs was substantially higher under WFTC. It also offered support for a wider age range, for children aged up to 15 (or 16 if disabled) rather than up to 11.[45] 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 costs.

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 arrangements for nonworking single mothers (although these women 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 older children only from June 2000 onward).[46] 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 approximately 3 percentage points. This average effect reflects an increasing impact over time, from 2.3 percentage points in 1999 to 3.7 percentage points in 2001 (these estimates are comparable to statistics drawn from FACS data reported in McKay [2002 and 2003]). At the same time, the corresponding childcare expenditures went up by about £16-17 per week in each of the three post-reform observations. As compared to all lone mothers, those with exactly one child aged 0-4 experienced an even higher increase in the probability of using paid childcare services in excess of 4 percentage points, compared to an increase of 3 percentage points for those with one child aged 5-10 (Table 6). Likewise, the WFTC effects on weekly childcare costs for those two types of women were similar or just slightly smaller than those reported in Table 5.

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[45] Under both WFTC and FC, however, only registered childcare (e.g., nurseries and registered childminders) counted towards eligible support. This, therefore, excludes informal childcare arrangements. Analysis based on the Families and Children Study (FACS) shows that only one quarter of the nonworking lone parents in 2001 were expecting to use formal childcare upon returning to work (McKay, 2003).

[46] Because the childcare information in the BHPS covers only ages 11 and 12 in the oldest group, we chose to estimate the WFTC effects on childcare use and costs only for the first two groups, ages 0-4 and 5-10 (Table 6).

Finally, there were no anticipation effects. The changes in lone mothers' labor market behavior observed before the implementation of the reform were not accompanied by the use of eligible childcare services. Lone mothers then might have used informal childcare arrangements not covered by FC, e.g., relatives, neighbours, and unregistered childminders (see Section VI). The timing of the increase in childcare use corresponds to the date at which the more generous childcare subsidies were actually implemented.

F. Marriage and Fertility Rates

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,[47] this literature is virtually nonexistent for Britain.

The seventh row in Table 5 and 6 presents causal effect estimates from linear probability models of transitions into partnership (marriage or cohabitation) for the sample of single childless women and lone mothers. For each woman, the dependent variable takes value zero if the woman remains single, and value one if she married (after that point her observations are censored). The eighth row in each table report estimates from models of transitions into birth for the subsample of lone mothers from the second year they were in the panel onwards. The estimates in Table 5 indicate that the implementation of WFTC led to a reduction in partnership rates.[48] Using single childless women as comparison group, the estimates imply that on average, with the WFTC reform in place, lone mothers were about 2 percentage points less likely to form a union. This effect is large, representing a 25-percent change with respect to the average annual (re)-partnership rate of 8.5 percent during the sample period (see

[47] See, for example, Schultz (1994), Moffitt (1998), Schoeni and Blank (2000), Bitler *et al*. (2002), Hotz and Scholz (2003) and references therein.

[48] This result is qualitatively similar to much of the available evidence for the United States (e.g., Bitler *et al*., 2002).

Table A1). There is a modest, though insignificant negative anticipation effect on partnership formation, and there is no indication that the negative effect declined over time, on the contrary it possibly increased. Most of that effect was driven by mothers of pre-school children (Table 6), who experienced a reduction in the chances of forming a partnership by 3.5 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 children in all age groups we found again stronger negative and significant effects ranging between 1.5 and 3 percentage points (not shown).

These changes were accompanied by a comparably large reduction in fertility rates among lone mothers. The post-reform risk of having an additional child decreased by 1.4 percentage points (which represents a 38-percent change over the average annual birth rate for lone mothers during the sample period, see Table A1), but this effect is statistically significant only at the 10 percent level. A significant effect however emerges for mothers with one child aged 0-4, who reduced the probability of another birth by about 2 percentage points. This was also the case when we considered the fertility response of mothers with two children both aged 0-4, or with one child aged 0-4 and the other aged 5-10.

G. Entry into Lone Motherhood

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.[49] The results in Table 5 (line 10) show little evidence of this unintended effect. In fact, the introduction of WFTC had the opposite impact, reducing the propensity of single

[49] 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.

childless women to form lone-mother households by about 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). This effect is however not statistically significant, nor are the other year-specific effects. Interestingly, these results are in line with those found 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.

VI. A Closer Look at Anticipation Effects

In the two previous sections we found the strongest anticipation effects to emerge for those lone parents who were expected to benefit the most, and who indeed showed the strongest responses to the WFTC reform, such as mothers of pre-school children. In addition, the greatest and most significant effects arose for employment and Income Support receipt, which suggest that lone mothers adjusted their labor market behavior (and thus their eligibility for IS) in anticipation of the WFTC reform. Other outcomes in 1998, instead, seem to have been unaffected (e.g., paid childcare use, FC/WFTC receipt, and entry into marriage). The lack of anticipation effects here may simply stem from ineligibility conditions (as is the case for paid childcare arrangements) or from 'time-to-build' investments (such as marriage and children).

The interpretation of these results as evidence of anticipation effects lines up well with the government's view of the WFTC at the time of its formal announcement in the Budget speech in March 1998 as a key reform 'to make work pay' for low-income families (H.M. Treasury, 1997 and 1998; Strickland, 1998). The Budget received wide media coverage, which is likely to have fostered a cultural climate that encouraged work and self-sufficiency among people in low-income families and with traditionally low labor market attachment, such as single parents.[50] Taken together, these various pieces of evidence seem to be consistent with an interpretation of our 1998 estimates representing anticipation effects. However, a greater confidence in this interpretation is perhaps gained after checking its robustness against alternative explanations, three of which are examined next.

[50] At that time, the government's dissemination effort was intense. This can be indirectly gauged by the number of press releases issued by the Treasury on 17 March 1998 (the day of the Budget speech). See ⟨http://archive.treasury.gov.uk/budget/1998/newsindx.htm⟩. A similar picture emerges from the post-Budget press releases by the Department of Social Security, which was then responsible for the administration of Family Credit (see ⟨http://www.dwp.gov.uk/mediacentre/dss⟩).

A. Other Policy Changes

The interpretation of the 1998 responses as anticipation effects relies on the absence of other policy changes in that year which could have differentially affected the outcomes of lone mothers compared to those of single childless women. Even though many in the latter group would have been eligible for other comparable programs that were introduced in 1997 and 1998, the New Deal for Lone Parents (NDLP) introduced in 1998 may have had such a differential impact. However, because the program only targeted lone parents whose youngest child was at least five years of age (the age cutoff was reduced to three only in April 2000), this could not explain the larger estimated behavioral response in 1998 of mothers with children between 0-4 years of age. In addition, by the end of 1998 only a small number of individuals had actually participated in the NDLP (Millar and Evans, 2003).

Another policy change that could have directly affected lone mothers' outcomes before the introduction of WFTC is the abolition of the higher lone-parent rate of Child Benefit in July 1998 (see Table 1 and the discussion in Section I). Although it cannot explain the WFTC effects in subsequent years (as, for example, it would fail to account for the differential responses based on child's age), this change instead might have triggered the estimated employment responses in 1998. More specifically, the repeal of the lone-parent rate, through a negative income effect, may have led to an increase in labor supply. To test this explanation, we estimated separate regressions for the probability of entering and the probability of staying in eligible employment distinguishing between lone mothers who previously received the lone-parent rate of CB and lone mother who did not. As the CB reform only affected the latter group, because the former were exempt from this policy change, we would expect differential employment effects for the two groups. However our results (not reported) do not confirm this expectation. Instead, they reveal that the 1998 effects shown in Table 4 are statistically identical for the two groups of mothers regardless of whether or not they received the more generous lone-parent rate. Similar employment transitions to those reported in Table 4 for the two groups of women also emerged by child's age and number. With a great deal of stability in women's receipt of the lone-parent rate of CB or One Parent Benefit over the entire sample period even after 1998, which indicate small compositional effects, the uniform responses across the two groups of single mothers suggest that the estimated 1998 employment behavior could not have been driven by the CB reform.

It is possible that the 1998 responses were not only in anticipation of the WFTC reform, but also in anticipation of the 1999 introduction of the minimum wage and the introduction of the 10p-income-tax band. Given the absence of significant behavioral responses by single women without children in 1998, this could only be the case if both 1999 reforms were expected to affect the two groups of women differently. There is however little support for this and in fact for any significant employment effects of such reforms in the British context (Connolly and Gregory, 2002; Stewart, 2004).[51]

B. Timing of the Government's Announcements and Labor Market Changes

As mentioned earlier, press coverage around the time of the Budget speech in March 1998 — when WFTC was formally announced — was considerable. For example, we counted 73 stories in The Times, The Daily Telegraph and BBC's Online News service that discussed the WFTC reform in 1998. Fifty-five of these were during the February-April period.[52] The existence and timing of anticipation effects is based on information diffusion, similar to the way that news announcements matter in financial markets (Andersen *et al*., 2003). We therefore examine whether or not our results match with the timing of that announcement. In particular, if the employment changes estimated for 1998 were concentrated after March of that year (seasonal effects apart), they could be seen as a behavioral adjustment in response to the announcement in 1998 of the 1999 policy reform. On the other hand, if such changes were equally spread over time before and after the Budget speech, that interpretation would be harder to defend. Figure 4 shows monthly rates in eligible employment for single women without children and for lone mothers between September 1997 and December 1998 (Panel (a)).[53] While the employment rate of single childless women increased by less than 1 percentage point over that period, the employment rate of

[51] Also note that the actual minimum wage levels, which were introduced in 1999, were only announced on the 25th of June 1998, while we find large employment responses already before that date (see below).

[52] Strömberg (2004) finds evidence which strongly supports the notion that mass media (radio in his case) have a sizeable impact on people's and politicians' behavior. Similarly, Heim (2004) in analyzing the effects of tax rebates on consumer spending finds large announcement effects.

[53] To construct such monthly figures we used the information that the BHPS collects on respondents' employment history over the period September 1st of the year prior to interview and the interview date.

single mothers increased by 3.5 percentage points (from 38 to 41.5 percent). Interestingly, 82 percent (about 2.9 percentage points) of that increase occurred during the April-July period. This result is consistent with the presence of anticipation effects.

Another way of documenting the importance of the WFTC announcement is to look at the distribution of dates (months) in which women who were in employment at the time of the 1998 interview started a job since the previous September. The distributions for single childless women and lone mothers, which include both job-to-job transitions and new labor market entries, are plotted in Figure 4, Panel (b). Of the 124 new jobs that lone mothers started between September 1997 and December 1998, 62 percent commenced after March 1998, whereas only 51 percent of the 186 new jobs by unmarried women without children began after the Budget speech.

C. Unpaid Childcare Use

An interesting feature of our results is that while the post-reform employment increases were accompanied by increases in paid childcare use, the relatively large increase in eligible employment in anticipation of the WFTC reform, in particular for those with young children, was not accompanied by a significant increase in paid childcare use. While the increase in paid childcare use corresponds well with the actual increase in childcare subsidies under WFTC, it raises the question of what childcare arrangements these women had that allowed them to increase their hours of work. The results suggest that single mothers who entered eligible employment in 1998 may have used unpaid childcare arrangements in that year in anticipation of using formal arrangements under WFTC, which offered a more generous childcare credit than Family Credit did (see Section VII). To check this hypothesis we estimated fixed-effects linear probability models of unpaid childcare use among working mothers, employing similar specifications to those used in Tables 5 and 6 for paid childcare. The results (not shown) indicate that there was indeed a temporary increase (albeit significant only at the 10 percent level) of about 2 percentage points in unpaid childcare use among all lone mothers in 1998. For mothers with one child aged 0-4, that increase was larger (approximately 3.3 percentage points) and significant. As we find no evidence of significant effects on unpaid childcare use in subsequent years, this temporary expansion lines up well with the post-reform increases in WFTC receipt, eligible employment, and formal childcare use.

When taken together, all these results — the fact that other policy changes may have played only a minor role at best, the link between labor market changes and timing of the reform announcement, and the switching from unpaid to paid childcare arrangements services before and after the WFTC reform took place — are consistent with our hypothesis that lone mothers did adjust their behavior in anticipation of the reform.

VII. What Caused the Large Effects for Mothers of Pre-school Children?

The results in Sections IV and V provide several clues for uncovering the main reasons underpinning the large employment responses to the introduction of WFTC. The strongest employment effects of the reform emerged for single mothers with young (pre-school) children. For this same group we also found the largest increases in WFTC receipt and childcare use. These findings point to two parameters of the reform, which appear to have played a role in generating those differential effects. First, as discussed in Section I, the reform substantially increased the credit received for younger children in the age group 0-10. Even though that increase was accompanied by a similar increase in the child-related allowances paid to nonworking single mothers through Income Support (see Table 1), the reform may have accelerated the decline in IS receipt among lone mothers, especially among those with pre-school children (Tables 5 and 6). Second, the estimates point to the importance of the childcare tax credit component.[54]

A. The Child Credit Component

To investigate whether this parameter played a specific role, we compare changes in labor supply from 1998 to 1999 between lone mothers who had a child aged 10 in 1998 and lone mothers who had a child aged 11 in that year (Sample 1). As the former group experienced a much larger increase in the child credit component of WFTC 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. The resulting difference-in-

[54] Clearly, other parameters responsible for the increased generosity of WFTC (such as the reduction of the taper rate or the increase in the applicable amount) may have had a part in explaining the employment impact of the reform. However, because they affected lone mothers uniformly regardless of their children's age, it is unlikely that they can explain the differential employment effects.

difference estimates are presented in Table 7 for both eligible employment and full-time employment. 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).[55] As there is no reason to expect a large difference in employment behavior between the two groups of mothers in absence of WFTC's differential child credit rules, a difference in employment behavior following the introduction of WFTC can be attributed to the differential child credit components.

However, the Wald estimates in Table 7 (column (v)) indicate that the child credit increase for young children does not play an important role in explaining the employment effects of the reform. Controlling for differences in individual characteristics (column (vi)) and increasing the two comparison samples to mothers with children aged 6-10 and 11-14 respectively (Sample 2) do not alter this result.

B. The Childcare Tax Credit Component

For this parameter, we first 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 previously working single mothers who did not use paid childcare started using paid childcare while working. Estimates from both analyses are reported in Table 8, where we show estimates obtained from specifications that allow for a pre-program trend but impose a constant treatment effect (column (i)) and estimates obtained from specifications that allow for year-specific treatment effects (column (ii)). They reveal that a significant fraction of those who entered eligible employment as a result of the reform did so by also choosing to use paid childcare. In fact the estimates in Tables 8 when compared to those in Table 4 suggest that, for mothers with one child aged 0-10, this work-childcare combination increased by about 3-4 percentage points (specification (i)), and accounted for almost half of the increase in the labor market entry rate. Similarly, we find that the reform led to an increase in the use of paid childcare amongst those who were already working 16 or more hours but not using paid childcare. In the case of mothers with one child aged 0-4 the rate of childcare use increased by 4.5 percentage points

[55] 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.

(column (i)). For both transition rates, the largest effects are consistently found among mothers with children in that age group, but — as it emerged when we analyzed childcare usage in isolation (Section V) — there is no evidence of anticipation effects.

Another way to illustrate the importance of the childcare subsidy in explaining the estimated employment responses is to relate it directly to the demand for paid childcare 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 six-year olds. Any difference in response behavior between the two groups could then be legitimately attributed to differential childcare needs. In Table 9 we present difference-in-difference estimates, which again have a simple RD interpretation. They indicate large remaining differences in lone mothers' employment responses, even after controlling for observable characteristics. Mothers of four-year olds in 1999 appeared to have increased their supply of eligible employment by about 3-4 percentage points more than mothers of six-year olds (column (vi)).

Table 10 further explores this point. The table shows least-squares estimates of the WFTC effect on eligible employment and full-time employment separately by child age for children between the ages of two and eight years.[56] The estimates come from two specifications, one that excludes and the other that includes anticipation effects (columns (i) and (ii) respectively). The table documents that the largest employment responses are found among mothers of three- and four-year olds, with no discernable effect for mothers of five and positive but smaller effects for mothers of children between the ages of six and eight. For example, after the introduction of WFTC, mothers of three- and four-year olds increased their eligible/full-time employment rates by 11-19 percentage points, while mothers of five- and six-year olds increase their labor supply by 1-4 percentage points. As the bottom two lines of the table indicate, the differences in employment responses between mother of four-

[56] Owing to small sample sizes, fixed-effects regressions were not estimated.

and five-year olds and four- and six-year olds are both statistically significant at the 5 percent level.

By and large, the estimates in Tables 8-10 provide strong evidence that the childcare subsidy component of WFTC played a key role in producing the estimated large employment effects for single women with young children. We instead do not find any labor supply response driven by the increase in the child credit component of WFTC. On the other hand, the simultaneous reduction in the taper rate and the increase in the child maintenance disregard are likely to have played an additional important role in the increase in the WFTC caseload (Section V) and the overall positive employment effect for all lone mothers. These separate components, however, were not assessed here, and deserve more attention in future work.

VII. 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 makes a number of improvements over past research. It identifies the effect of the reform through comparisons of changes in behavior for lone mothers and single women without children by both exploiting the panel nature of our data and accounting for the possible presence of anticipation effects. We also consider effects on a wider set of socioeconomic outcomes, some of which may have been unintended. Finally, we use some aspects of the reform design and the panel nature of the data to identify which parameters of the reform were more likely to explain its estimated effects.

We stress six 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 7 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 (30 or more hours per week). Second, this large employment 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 12 percentage points. We instead find no effects for mothers with multiple older children. Fourth, there are important (and perhaps unintended) effects on lone mothers' behaviors other than on employment. In particular, there is evidence of significant reductions in single mothers' subsequent fertility and in the rate at which they married. 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. More than 50 percent of the increased entry rate in eligible employment was attributable to lone mothers who also chose paid childcare arrangements, and the effect was stronger for mothers with pre-school aged children. Similarly, among single mothers who continued to be in employment, 50 percent of their greater post-reform labor market attachment is observed in conjunction with paid childcare services. Sixth, we find relatively large anticipation effects, especially for the responses in employment behavior and Income Support receipt. Ignoring such effects when using evaluation methods that rely on before-after comparisons could lead to significant biases. A direct implementation of a DD or DDD approach using 1999 as the intervention year, would have led to a downward bias of 2.5 percentage points (36 percent) in the overall employment effect. This underlines the importance of considering their presence in other models of program evaluation.

The results that WFTC, and especially its childcare tax credit component, played a key role in the employment increases for lone mothers suggest that in-work benefit policies are effective in encouraging work among lone mothers. In fact, with the most powerful effects emerging for women with children under five, these results emphasize how crucial the focus on incentives can be for designing successful welfare policies (Moffitt, 2003b). 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. 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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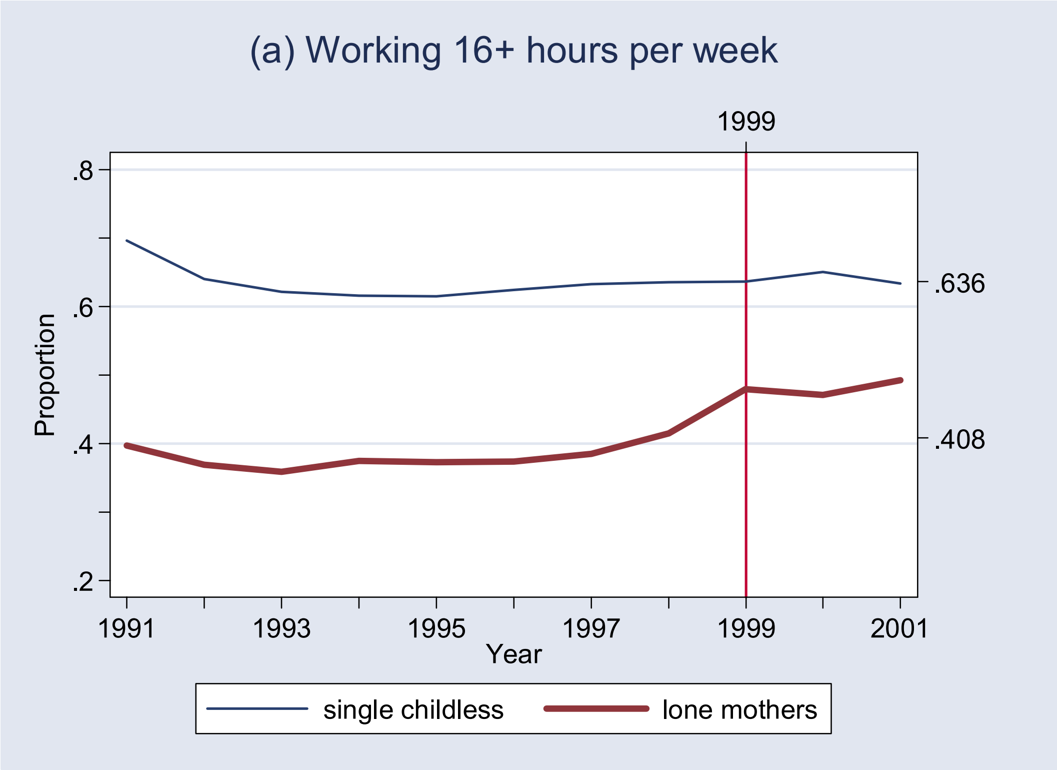
Figure 1. Working 16 or More Hours per Week – Single Childless Women and Lone Mothers



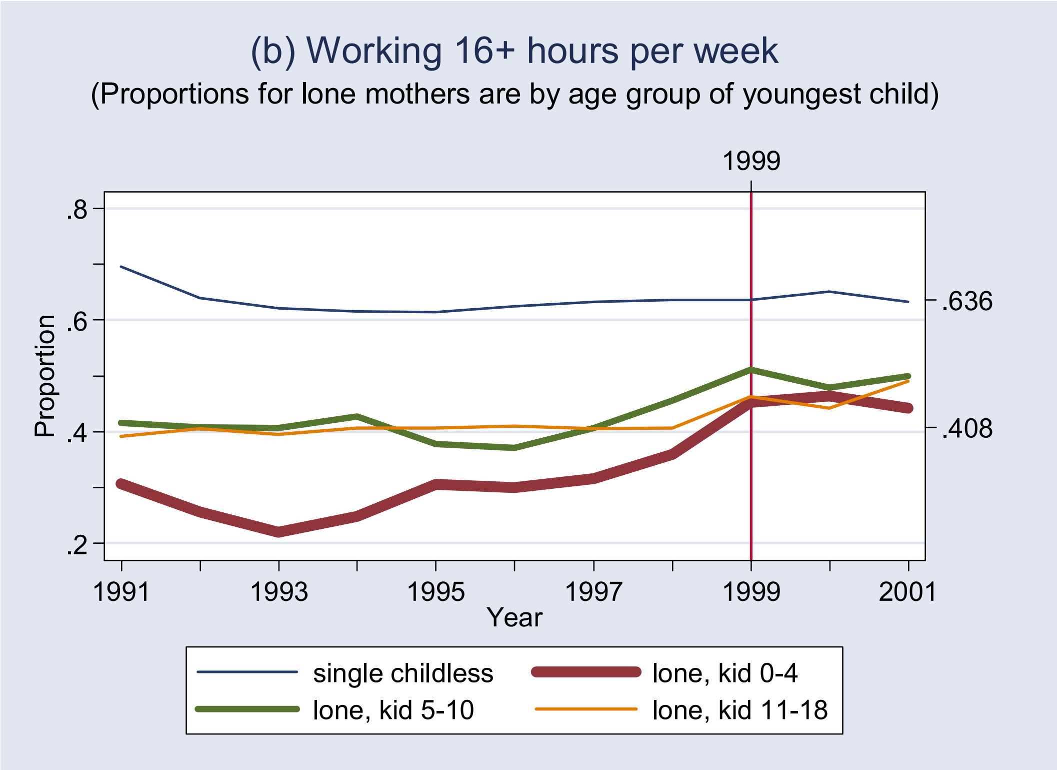
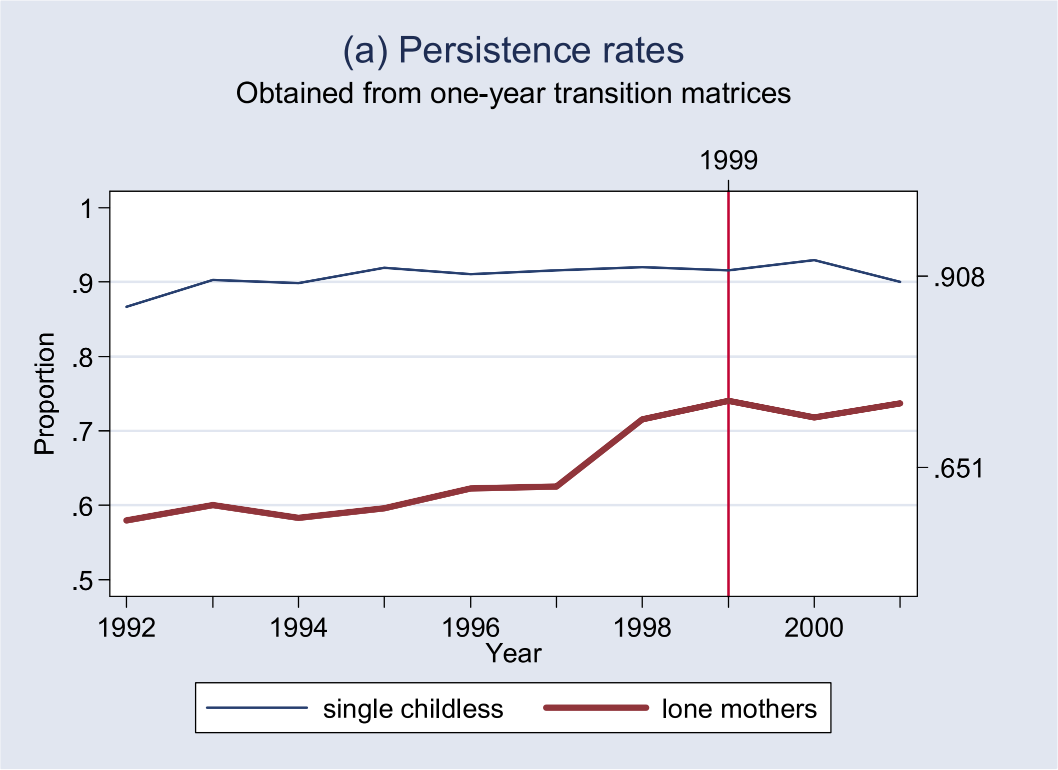


Figure 2. One-Year Transition Rates in the Probability of Working 16 or More Hours per Week – Single Childless Women and Lone Mothers



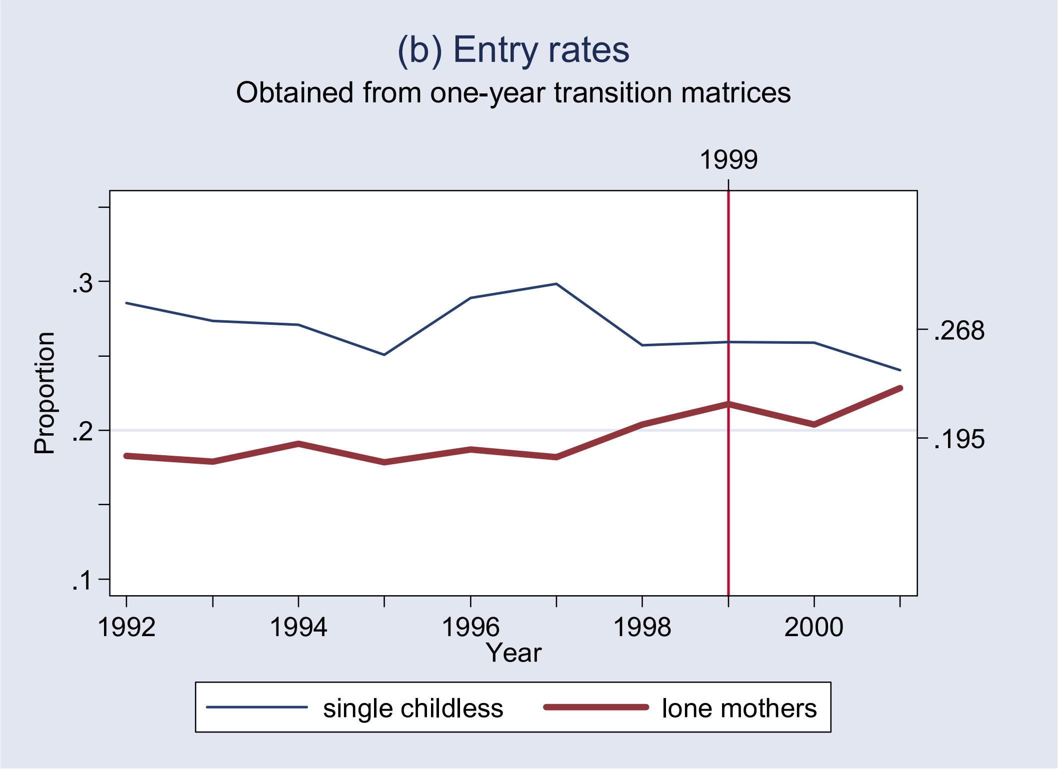
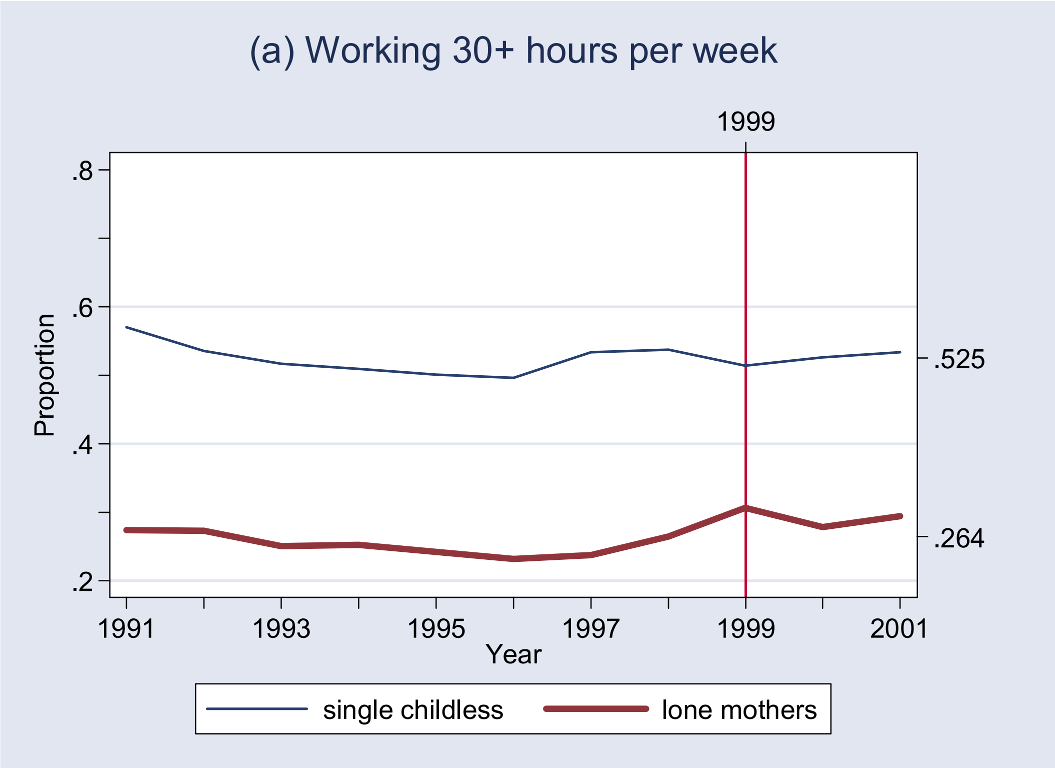
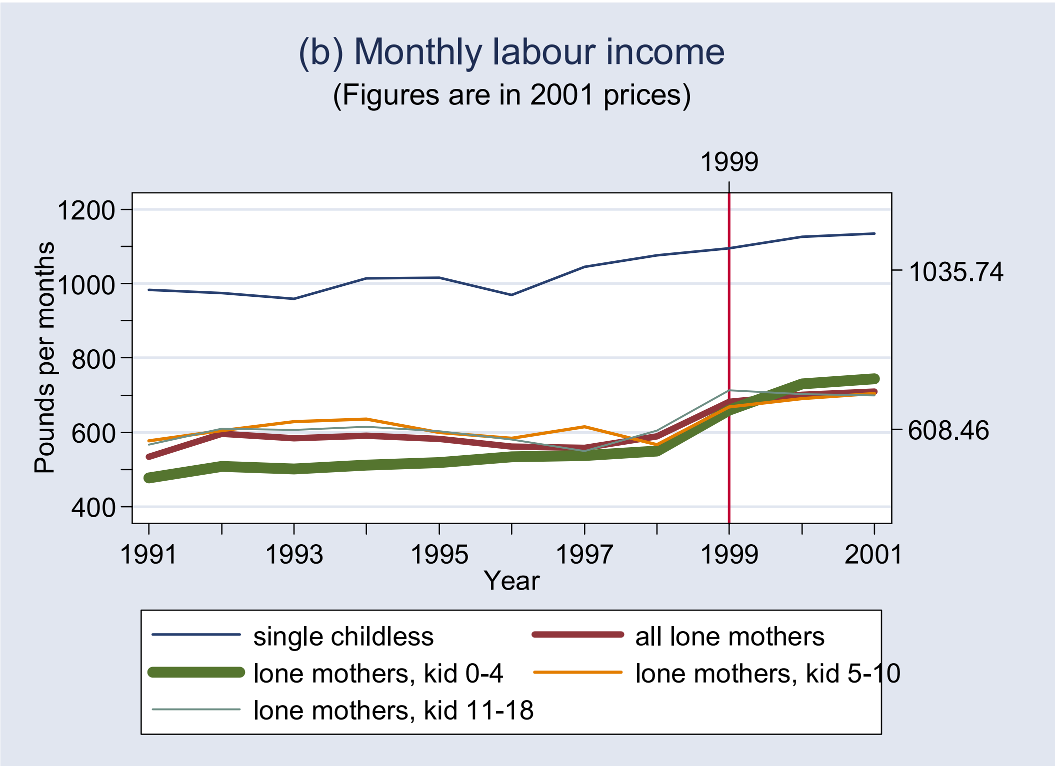
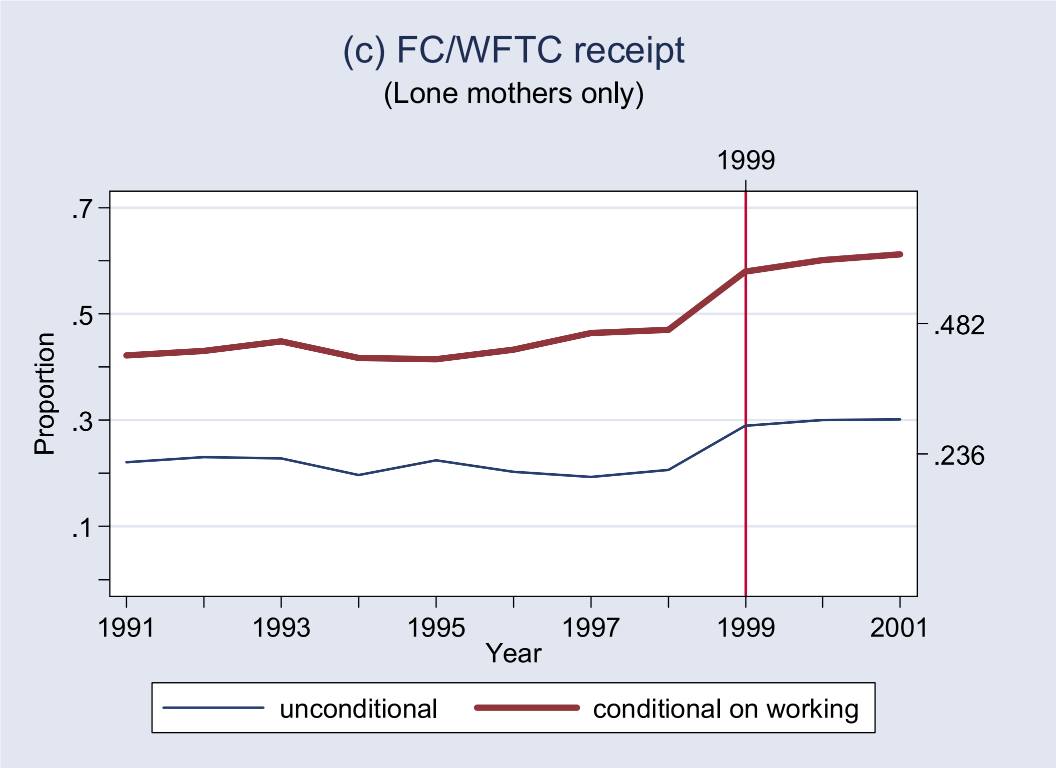
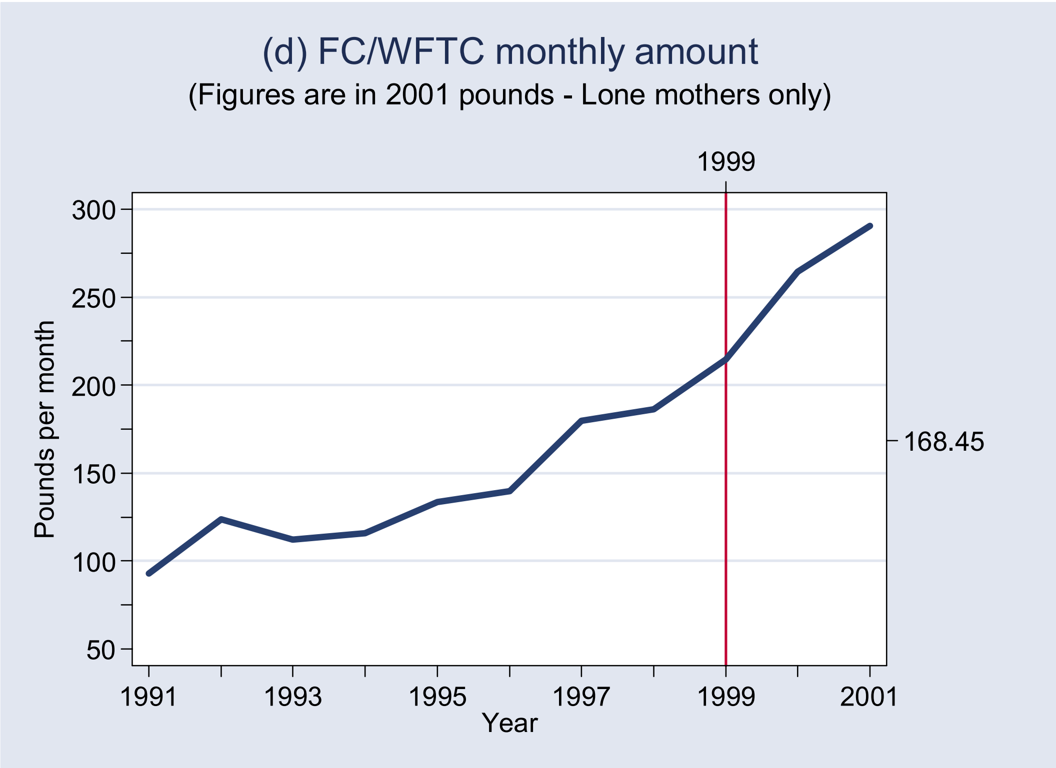


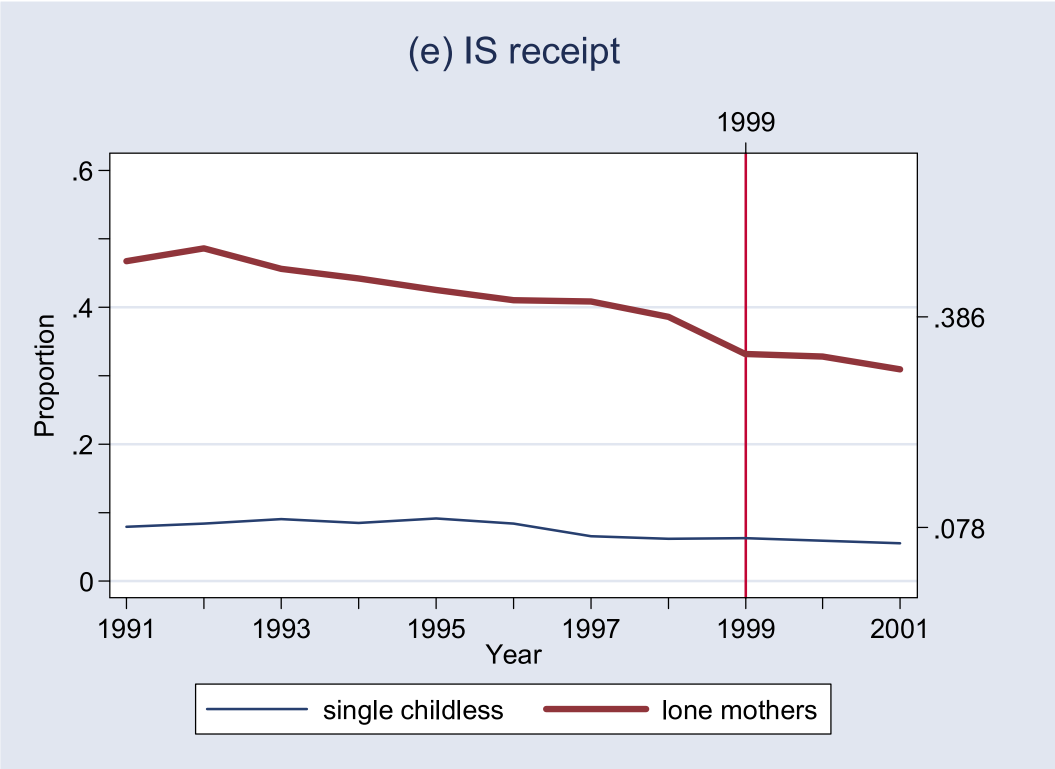
Figure 3. Other Outcomes for Lone Mothers and Single Childless Women



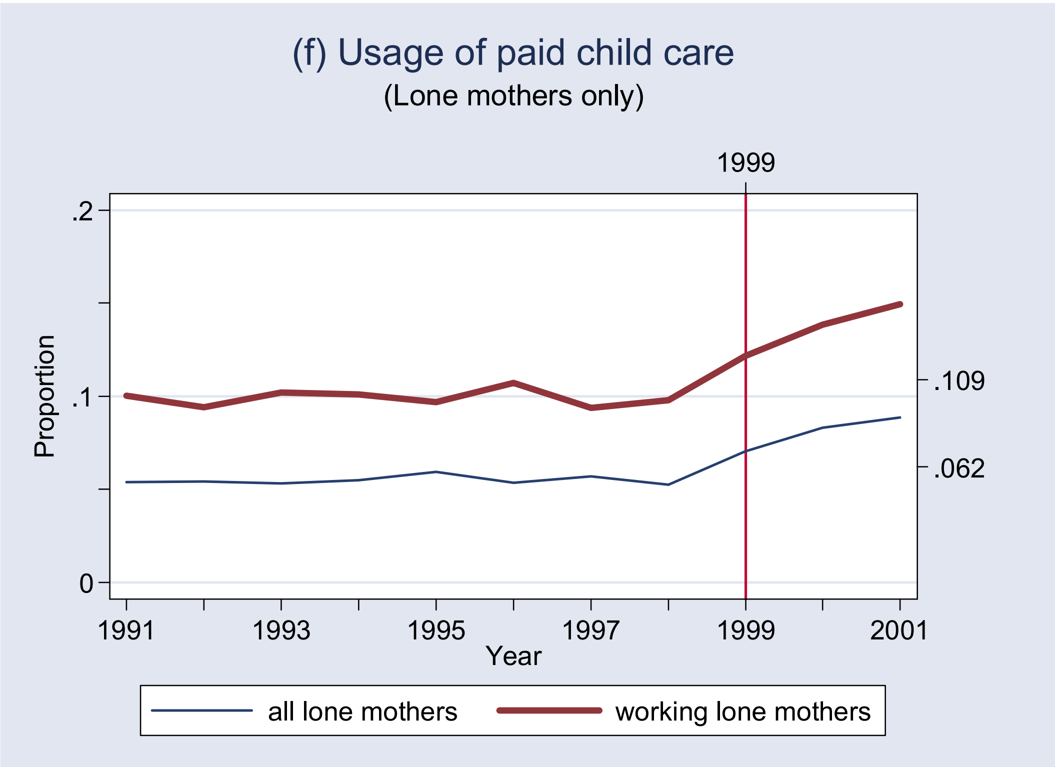



(c) FC/WFTC receipt
(Lone mothers only)
1999
.7
.5
.3
.1
Proportion
.482
.236
1991
1993
1995
1997
1999
2001
Year
unconditional
conditional on working

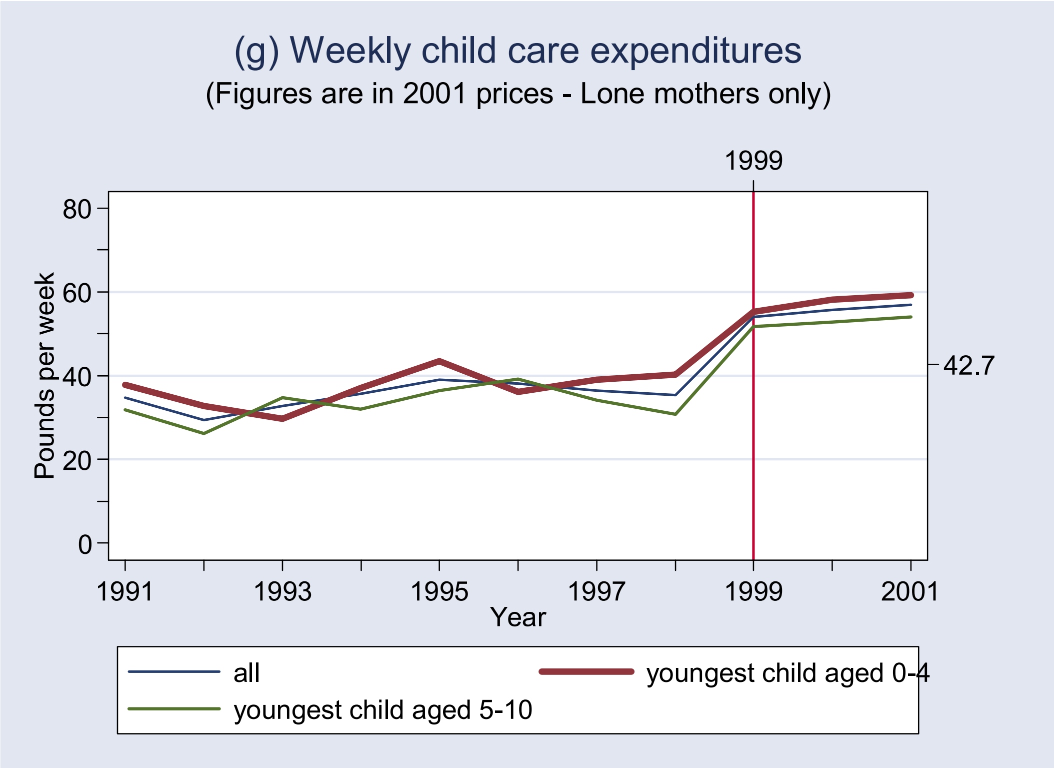

(d) FC/WFTC monthly amount
(Figures are in 2001 pounds - Lone mothers only)
1999
300
250
200
150
100
50
Pounds per month
168.45
1991
1993
1995
1997
1999
2001
Year

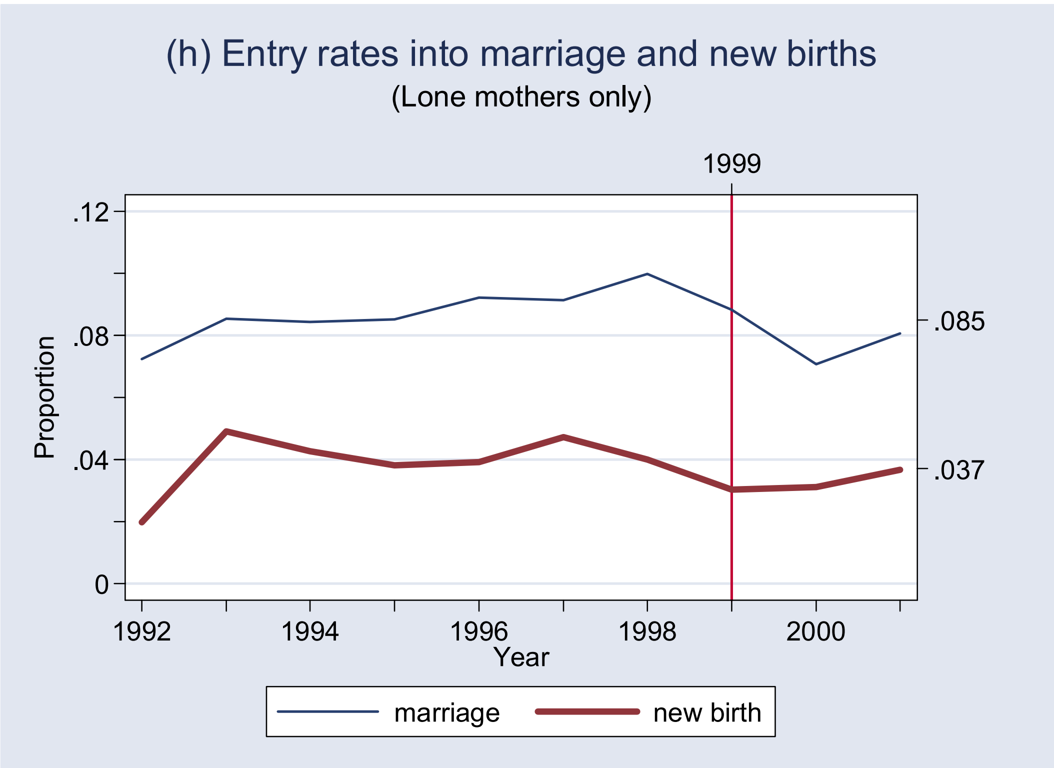


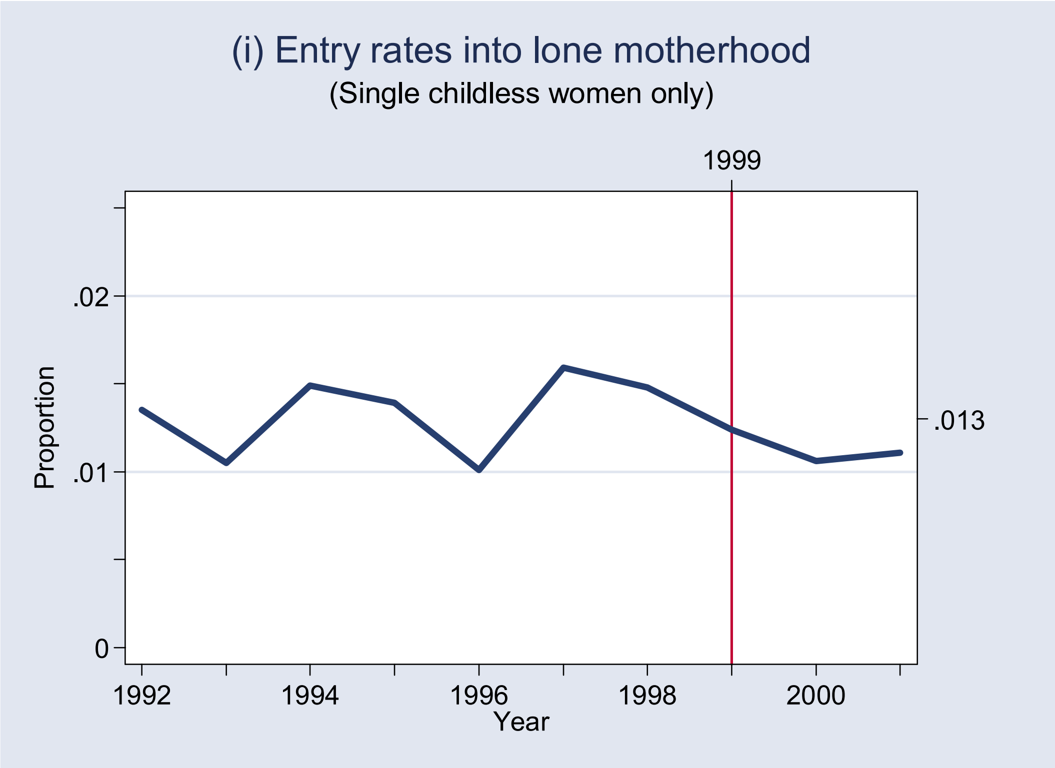
(e) IS receipt
1999
.6
.4
.2
0
.386
.078
Proportion
1991
1993
1995
1997
1999
2001
Year
single childless
lone mothers



(f) Usage of paid child care
(Lone mothers only)
1999
.2
.1
0
.109
.062
Proportion
1991
1993
1995
1997
1999
2001
Year
all lone mothers
working lone mothers

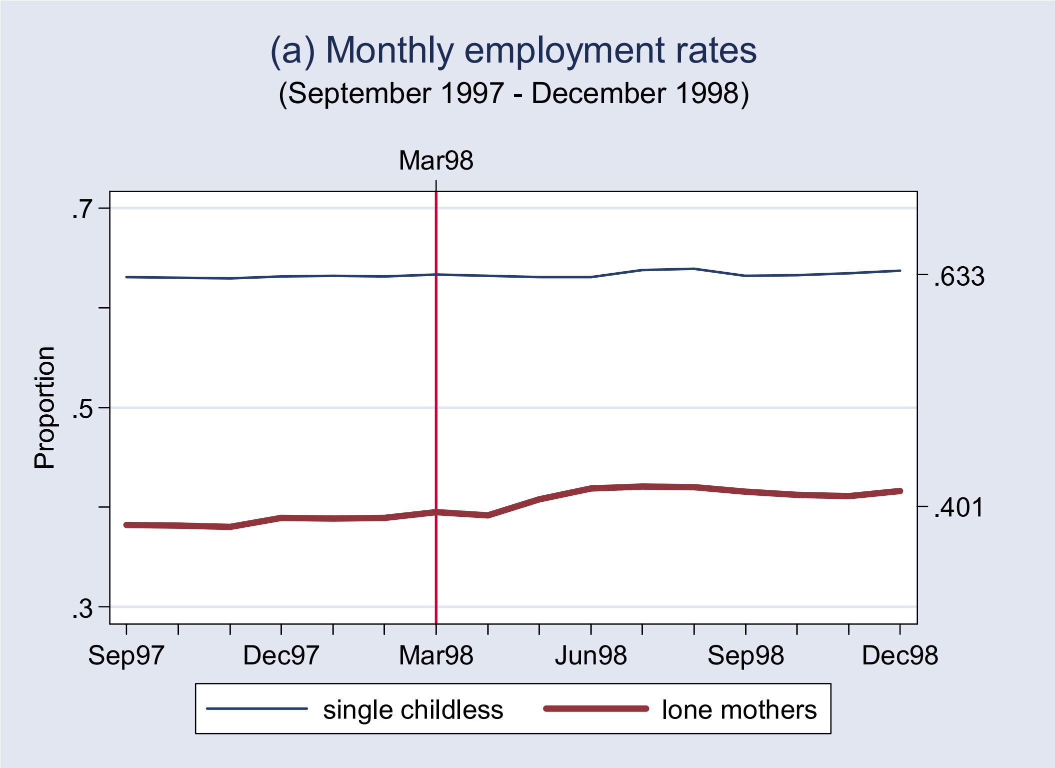

(g) Weekly child care expenditures
(Figures are in 2001 prices - Lone mothers only)
1999
80
60
40
20
0
Pounds per week
42.7
1991
1993
1995
1997
1999
2001
Year
all
youngest child aged 0-4
youngest child aged 5-10


(h) Entry rates into marriage and new births
(Lone mothers only)
1999
.12
.08
.04
0
Proportion
.085
.037
1992
1994
1996
1998
2000
Year
marriage
new birth



(i) Entry rates into lone motherhood
(Single childless women only)
1999
.02
.01
0
.013
Proportion
1992
1994
1996
1998
2000
Year

Figure 4. Monthly Employment Rates and Timing of New Jobs Between 1997 and 1998 – Single Childless Women and Lone Mothers



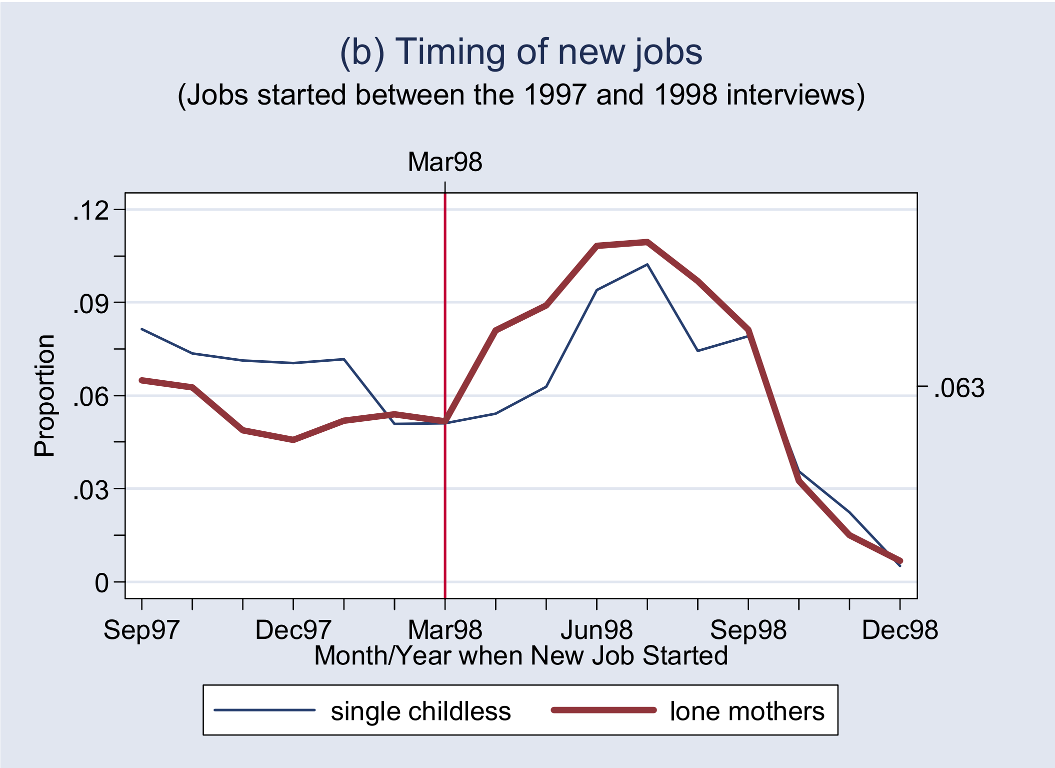


Table 1. Details on some of the major welfare benefits and credits supporting families with children over the sample period

	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
1. Family Credit/Working Family Tax Credit[a]											
Basic rate	£ 38.30	£ 41.00	£ 42.50	£ 44.30	£ 45.10	£ 46.45	£ 47.65	£ 48.80	£ 52.30	£ 53.15	£ 59.00
Additional credit for working 30+ hours a week	None	None	None	None	£ 10.00	£ 10.30	£ 10.55	£ 10.80	£ 11.15	£ 11.25	£ 11.45
Credit for child aged:											
0-10	£ 9.70	£ 10.40	£ 10.75	£ 11.20	£ 11.40	£ 11.75	£. 12.05	£ 14.85	£ 19.85	£. 21.25	£ 26.00
11-15	£ 16.10	£ 17.25	£ 17.85	£ 18.55	£18.90	£ 19.45	£ 19.95	£ 20.45	£ 20.90	£. 21.25	£ 26.00
16-17	£ 20.05	£ 21.45	£ 22.20	£ 23.05	£ 23.45	£ 24.15	£ 24.80	£ 25.40	£ 25.95	£ 26.35	£ 26.75
18	£ 27.95	£ 29.90	£ 31.00	£ 32.20	£ 32.80	£ 33.80	£ 34.70	£ 25.40	£ 25.95	£ 26.35	£ 26.75
Taper rate; threshold	70%; £62.25	70%; £66.60	70%; £ 69.00	70%; £71.70	70%; £73.00	70%; £75.20	70%; £77.15	70%; £79.00	55%; £90.00	55%; £91.45	55%; £92.90
Childcare tax credit	Costs are deducted from earnings	Costs are deducted from earnings	Costs are deducted from earnings	Costs are deducted from earnings	Costs are deducted from earnings	Costs are deducted from earnings	Costs are deducted from earnings	Costs are deducted from earnings	70% of up to: £100 (1 kid) £150 (2+ kids)	70% of up to: £ 100 (1 kid) £ 150 (2+ kids)	70% of up to: £ 135 (1 kid) £ 200 (2+ kids)
Disregard on child maintenance	None	None	£ 15.00	£ 15.00	£ 15.00	£ 15.00	£ 15.00	£ 15.00	Full	Full	Full
2. Income Support [b]											
Lone parent											
Basic rate (under 18)	£ 23.65	£ 25.55	£ 26.45	£ 27.50	£ 28.00	£ 28.85	£ 29.60	£ 30.30	£ 30.95	£ 31.45	£ 31.95
Aged 18 or over	£ 39.65	£ 42.45	£ 44.00	£ 45.70	£ 46.50	£ 47.90	£ 49.15	£ 50.35	£ 51.40	£ 52.20	£ 53.05
Family premium	£ 4.45	£ 4.75	£ 4.90	£ 5.10	£ 5.20	£ 5.20	£ 15.75	£ 15.75	£ 15.75	£ 15.90	£ 15.90
Couple											
Both under 18	£ 47.30	£ 50.60	£ 52.40	£ 54.55	£ 55.55	£ 57.20	£ 58.70	£ 60.10	£ 61.35	£ 62.35	£ 63.35
One/Both over 18	£ 62.25	£ 66.60	£ 69.00	£ 71.70	£ 73.00	£ 75.20	£ 77.15	£ 79.00	£ 80.65	£ 81.95	£ 83.25
Family premium	£ 8.70	£ 9.30	£ 9.65	£ 10.05	£ 10.25	£ 10.55	£ 10.80	£ 11.05	£ 13.90	£ 14.25	£ 14.50

Dependent children											
Under 11	£ 13.60	£ 14.55	£ 15.05	£ 15.65	£ 15.95	£ 16.45	£ 16.90	£ 19.80	£ 24.90	£ 30.95	£ 32.45
11-15	£ 20.00	£ 21.40	£ 22.15	£ 23.00	£ 23.40	£ 24.10	£ 24.75	£ 25.35	£ 25.90	£ 30.95	£ 32.45
16-17	£ 23.90	£ 25.55	£ 26.45	£ 27.50	£ 28.00	£ 28.85	£ 29.60	£ 30.30	£ 30.95	£ 31.75	£ 33.75
18	£ 31.40	£ 33.60	£ 34.80	£ 36.15	£ 36.80	£ 37.90	£ 29.60	£ 30.30	£ 30.95	£ 31.75	£ 33.75
3. Child Benefit/One Parent Benefit[c]											
Eldest eligible child:											
One Parent Benefit / Lone parent rate	£ 5.60	£ 5.85	£ 6.05	£ 6.15	£ 6.30	£ 6.30	£ 17.10	(£ 17.10)	(£ 17.10)	(£ 17.55)	(£ 17.55)
Standard rate	£ 9.25	£ 9.65	£ 10.00	£ 10.20	£ 10.40	£ 10.80	£ 11.05	£ 11.45	£ 14.40	£ 15.00	£ 15.50
Other children (each)	£ 7.50	£ 7.80	£ 8.10	£ 8.25	£ 8.45	£ 8.80	£ 9.00	£ 9.30	£ 9.60	£ 10.00	£ 10.35
4. Statutory Maternity Pay and Maternity Allowance[d]											
SMP higher rate	None	None	None	90% of average weekly earnings (first 6 weeks)	90% of average weekly earnings (first 6 weeks)	90% of average weekly earnings (first 6 weeks)	90% of average weekly earnings (first 6 weeks)	90% of average weekly earnings (first 6 weeks)	90% of average weekly earnings (first 6 weeks)	90% of average weekly earnings (first 6 weeks)	90% of average weekly earnings (first 6 weeks)
SMP lower rate	£ 44.50 a week (up to 18 weeks)	£ 46.30 a week (up to 18 weeks)	£ 47.95 a week (up to 18 week)	£ 48.80 a week (up to 18 weeks)	£ 52.50 a week (up to 18 weeks)	£ 54.55 a week (up to 18 weeks)	£ 55.70 a week (up to 18 weeks)	£ 57.70 a week (up to 18 weeks)	£ 59.55 a week (last 12 weeks)	£ 60.20 a week (last 12 weeks)	£ 62.20 a week (last 12 weeks)
MA higher rate	£ 40.60	£ 42.25	£ 43.75	£ 44.55	£ 52.50	£ 54.55	£ 55.70	£ 57.70	£ 59.55	£ 60.20	
MA lower rate	£ 40.60	£ 42.25	£ 43.75	£ 44.55	£ 45.55	£ 47.35	£ 48.35	£ 50.10	£ 51.70	£ 52.25	
MA standard rate											£ 62.20 up to 18 weeks (to women with average earnings above a pre-specified limit)
MA variable rate											90% of average weekly earnings if average weekly earnings are less than a

											pre-specified limit
5. Social fund payments[e]											
Maternity grant	£ 100	£ 100	£ 100	£ 100	£ 100	£ 100	£ 100	£ 100	£ 100	£ 200	£ 300
6. Housing Benefit[f]											
Lone parent premium	£ 10.05	£ 10.60	£ 10.95	£ 11.25	£ 11.50	£ 11.50	£ 22.05	£ 22.05	£ 22.05	£ 22.20	£ 22.20
Fuel charges*	£ 10.70	£ 11.40	£ 11.40	£ 11.40	£ 12.15	£ 12.35	£ 12.35	£ 12.35	£ 12.35	£ 12.35	£ 12.35
Deductions for non-dependants:											
Maximum*	£ 13.50	£ 18.00	£ 21.00	£ 25.00	£ 30.00	£ 32.00	£ 39.00	£ 45.00	£ 46.35	£ 47.75	£ 47.75
Minimum*	£ 5.70	£ 4.00	£ 4.00	£ 5.00	£ 5.00	£ 6.00	£ 7.00	£ 7.00	£ 7.20	£ 7.40	£ 7.40

Source: Child Poverty Action Group (various issues).

[a] Eligibility (means-tested): Must work 16 hours or more a week (this minimum was 24 hours or more up to 1991); have at least one dependent child (i.e., under 16, or aged under 19 an in full-time education); family income is sufficiently low; savings and capital should be worth less than £8,000 (this sum has not changed over the whole sample period); couples need to claim jointly; need not be married. Income is defined as net income (i.e., as income after tax and national insurance), and some benefit income is excluded from the definition of income (e.g., Child Benefit, Statutory Maternity Pay, Maintenance payments, Housing Benefit and Council Tax Benefit awards. Notice, however, that Housing Benefit and Council Tax Benefit count as income in the computation of their awards). The credit is paid biweekly or monthly and over a period of 26 weeks by the Inland Revenue directly to the applicant through the wage packet by the employer (WFTC only) or by automatic credit transfer to a bank account, direct debit, order book or girocheque.

[b] Eligibility (means-tested): Benefit for people with low income. Income support is not paid to unemployed people available for and actively seeking work who may claim jobseeker's allowance; and it is not paid to people working 16 or more hours per week who may claim Family Credit or Working Family Tax Credit. Must be aged 16 or more, not studying full-time, not working 16 or more hours per week (including partner); family income is sufficiently low; savings and capital should be worth less than £8,000 (this sum has not changed over the whole sample period). People can claim Income Support if they satisfy the rules above and fit into one of the following groups: aged 60 or more; incapable of work because of illness or disability; lone/foster parent of a child under 16; pregnant and unable to work; looking after partner. The amount of Income Support is paid weekly by girocheque, benefit order book or directly into a bank or account, and is given by the difference between the 'applicable amount' (which measures needs, and consists of personal and children allowances, family premiums and housing costs, as given in this table) and income (from other benefits, part-time earnings and maintenance).

[c] Eligibility (non-means-tested): Must be responsible for at least one child (i.e., a person under 16 or a person aged 16 or over but under 19 and receiving full-time non-advanced education). The lone parent rate of Child Benefit, which replaced the former One Parent Benefit in April 1997, was abolished from 6 July 1998. Some lone parents may still be able to receive the lone parent rate of Child Benefit if some conditions are satisfied (e.g., parent was receiving the lone parent rate of Child Benefit on 5 July 1998 and has continued to qualify for it since that time).

[d] Eligibility (non-means-tested): Woman must be pregnant and within the last 11 weeks before the expected week of childbirth or have recently given birth. In the case of Statutory Maternity Pay (SMP) the woman must also be an employee for a continuous period of 26 weeks (or unemployed and become unemployed after week 26 of pregnancy), have stopped work for the employer paying SMP and must not work for other employers after the birth. In the case of Maternity Allowance (MA), the woman must not be entitled to SMP, and must be self-employed, or must not be currently working and have been employed and/or self-employed for a minimum number of week over a pre-specified period preceding the expected week of childbirth, or must have changed jobs during pregnancy. Even the higher rates of SMP or MA represent the minimum amount of pay that the law requires employers to pay during maternity leaves. Specific groups of employees (or trade unions) may have negotiated higher amounts (for longer periods of time). Before the 1993-94 reforms, coverage depended on length of service and was not universal. Eligibility and amounts paid were also different (e.g., two years of continuous employment at more than 16 hours per week with the same employer, or five years at more than 8 hours per week with the same employer). For some workers, the higher rate of SMP could have already been set at 90% of the average weekly pay. See Waldfogel (1998).

[e] Eligibility (means-tested): Must receive a 'qualifying benefit' (i.e., IS or FC/WFTC) on the date of claim; claimant (or her partner) must be pregnant or have given birth in the last three months; family capital must not exceed £500 (this sum has not changed over the whole sample period). If family capital exceeds £500, the payment is reduced by £1 for every £1 of capital inc excess of £500 (i.e., a capital in excess of £600 implies no payment). The amount in the table refers to the payment paid for each child. From June 2000, this benefit has been known as 'Sure Start maternity grant'.

[f] Eligibility (means-tested): Must have low income; savings and capital should be worth less than £16,000 (this sum has not changed over the whole sample period); must pay rent and the payment can be met by Housing Benefit (HB), that is, HB can meet payments that are not normally treated as rent, such as payments as a licensee and payments for hostel accommodation. The amount of HB mainly depends on: the 'applicable amount', which represents needs and is equivalent to the personal allowances and family premiums of IS shown above (the only relevant exception is the lone parent premium shown here); the maximum HB (i.e., eligible rent minus deductions); and family income. Individuals on IS are entitled maximum HB. Individuals who are not on IS receive maximum HB if their income is less than their applicable amount. If their income is greater than the applicable amount, they are entitled to the maximum HB minus 65% of difference between income and applicable amount. The benefit can be obtained even if the claimant does not work full-time and can be paid in conjunction with other social security benefits, or by itself. Paid by local authorities (although it is a national scheme).

* Figures (which are the sum of flat-rate deductions for heating, hot water, lighting and cooking) refer to the case when family occupies more than one room. The 'maximum' deduction refers to non-dependant aged 18 or over and in full-time work with weekly income above a specified threshold; the 'minimum' deduction refers to non-dependants with income below another lower level of weekly income.

Table 2. The Effect of Welfare Reforms on Lone Mothers' Probability of Working 16 or More Hours per Week (N=14,357)

Type of estimation and model	(i) $[\alpha_{31} = \alpha_{32} = 0;$ $\beta_{1999} = \beta_{2000} = \beta_{2001} = \beta]$		(ii) $[\beta_{1999} = \beta_{2000} = \beta_{2001} = \beta]$		(iii)			
	β	β_{1998}	β	β_{1998}	β_{1999}	β_{2000}	β_{2001}	β_{1998}
Level estimates ($\theta_1 = 0$)								
Basic model	**0.053**		**0.059**		**0.088**	**0.031**	**0.060**	
	(0.019)		(0.022)		(0.023)	(0.016)	(0.027)	
Basic model plus	**0.082**	**0.035**	**0.085**	**0.049**	**0.109**	**0.066**	**0.085**	**0.044**
1998 time dummy	(0.019)	(0.022)	(0.028)	(0.024)	(0.029)	(0.032)	(0.041)	(0.024)
Fixed-effects estimates								
Basic model	**0.053**		**0.047**		**0.071**	**0.020**	**0.042**	
	(0.014)		(0.018)		(0.019)	(0.019)	(0.024)	
Basic model plus	**0.067**	**0.040**	**0.073**	**0.044**	**0.093**	**0.047**	**0.069**	**0.040**
1998 time dummy	(0.016)	(0.018)	(0.021)	(0.020)	(0.022)	(0.024)	(0.027)	(0.019)

Notes: Standard errors are shown in parentheses. Estimates (in bold) are obtained from linear probability models on the sample of single childless women and lone mothers. See equation (7) for details. The vector **X** includes: 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 age and number of children by age group, age and the educational group dummies, and education and number of children by age group. N = number of person-wave observations.

Table 3. The Effect of Welfare Reforms on Lone Mothers' Probability of Working 16 or More Hours per Week by Number and Age Group of Children – Fixed-Effects Estimates (N=14,357)

Number and age group of children	(i) $[\alpha_{31} = \alpha_{32} = 0;$ $\beta^{(j)}_{1999} = \beta^{(j)}_{2000} = \beta^{(j)}_{2001} = \beta^{(j)}]$		(ii) $[\beta^{(j)}_{1999} = \beta^{(j)}_{2000} = \beta^{(j)}_{2001} = \beta^{(j)}]$		(iii)			
	$\beta^{(j)}$	$\beta^{(j)}_{1998}$	$\beta^{(j)}$	$\beta^{(j)}_{1998}$	$\beta^{(j)}_{1999}$	$\beta^{(j)}_{2000}$	$\beta^{(j)}_{2001}$	$\beta^{(j)}_{1998}$
One child:								
Aged 0-4	**0.110**	**0.061**	**0.117**	**0.062**	**0.132**	**0.074**	**0.102**	**0.060**
	(0.025)	(0.019)	(0.029)	(0.020)	(0.033)	(0.030)	(0.027)	(0.021)
Aged 5-10	**0.083**	**0.048**	**0.085**	**0.048**	**0.121**	**0.044**	**0.052**	**0.049**
	(0.024)	(0.029)	(0.026)	(0.030)	(0.035)	(0.018)	(0.023)	(0.028)
Aged 11-18	**0.044**	**0.014**	**0.046**	**0.019**	**0.076**	**0.035**	**0.058**	**0.013**
	(0.020)	(0.017)	(0.020)	(0.017)	(0.024)	(0.013)	(0.020)	(0.017)
Two children:								
Both aged 0-4	**0.048**	**0.018**	**0.046**	**0.019**	**0.057**	**0.046**	**0.012**	**0.020**
	(0.056)	(0.047)	(0.043)	(0.044)	(0.032)	(0.044)	(0.041)	(0.042)
Both aged 5-10	**0.031**	**0.017**	**0.020**	**0.016**	**0.039**	**0.013**	**0.022**	**0.017**
	(0.024)	(0.039)	(0.025)	(0.036)	(0.027)	(0.029)	(0.021)	(0.034)
Both aged 11-18	**0.066**	**0.014**	**0.062**	**0.017**	**0.070**	**0.054**	**0.059**	**0.018**
	(0.033)	(0.018)	(0.034)	(0.017)	(0.028)	(0.041)	(0.036)	(0.019)
One aged 0-4 and	**0.039**	**0.024**	**0.043**	**0.022**	**0.048**	**0.041**	**0.026**	**0.024**
one aged 5-10	(0.025)	(0.033)	(0.023)	(0.034)	(0.020)	(0.026)	(0.029)	(0.037)
One aged 0-4 and	**0.028**	**0.015**	**0.032**	**0.014**	**0.035**	**0.030**	**0.026**	**0.015**
one aged 11-18	(0.038)	(0.032)	(0.030)	(0.031)	(0.029)	(0.033)	(0.031)	(0.030)
One aged 5-10 and	**0.040**	**0.021**	**0.037**	**0.020**	**0.044**	**0.032**	**0.030**	**0.022**
one aged 11-18	(0.036)	(0.033)	(0.035)	(0.029)	(0.028)	(0.036)	(0.043)	(0.034)

Notes: Standard errors are shown in parentheses. Estimates (in bold) are obtained from linear probability models on the sample of single childless women and lone mothers. See equation (8) for details. Other definitions and (time-varying) variables used in the estimation are specified in the note to Table 2.

Table 4. The Effect of Welfare Reforms on Lone Mothers' Transition Probabilities of Working 16 or More Hours per Week

Type of transition and model	(i) $[\beta^{(j)}_{1999} = \beta^{(j)}_{2000} = \beta^{(j)}_{2001} = \beta^{(j)}]$		(ii)			
	β	β_{1998}	β_{1999}	β_{2000}	β_{2001}	β_{1998}
A. Persistence probability[a] (N=6,123)						
a. No distinction by child's age group						
Basic model plus 1998 time dummy	**0.068**	**0.037**	**0.083**	**0.021**	**0.024**	**0.033**
	(0.027)	(0.019)	(0.036)	(0.046)	(0.053)	(0.037)
b. Distinguishing by child's age group						
Basic model plus 1998 time dummy:						
One child aged 0-4	**0.122**	**0.016**	**0.138**	**0.087**	**0.052**	**0.027**
	(0.054)	(0.082)	(0.056)	(0.041)	(0.065)	(0.082)
One child aged 5-10	**0.067**	**0.079**	**0.093**	**0.023**	**-0.011**	**0.069**
	(0.038)	(0.033)	(0.041)	(0.051)	(0.054)	(0.033)
One child aged 11-18	**0.001**	**-0.043**	**0.036**	**-0.018**	**-0.027**	**-0.042**
	(0.040)	(0.033)	(0.040)	(0.046)	(0.053)	(0.034)
B. Entry probability[b] (N=5,114)						
a. No distinction by child's age group						
Basic model plus 1998 time dummy	**0.058**	**0.032**	**0.064**	**-0.022**	**0.043**	**0.036**
	(0.022)	(0.020)	(0.028)	(0.030)	(0.023)	(0.021)
b. Distinguishing by child's age group						
Basic model plus 1998 time dummy:						
One child aged 0-4	**0.076**	**0.062**	**0.112**	**0.034**	**0.056**	**0.060**
	(0.033)	(0.032)	(0.034)	(0.049)	(0.030)	(0.031)
One child aged 5-10	**0.045**	**0.031**	**0.055**	**-0.012**	**0.029**	**0.028**
	(0.032)	(0.029)	(0.031)	(0.047)	(0.053)	(0.029)
One child aged 11-18	**0.040**	**-0.038**	**0.038**	**-0.019**	**-0.012**	**-0.036**
	(0.029)	(0.030)	(0.040)	(0.043)	(0.048)	(0.027)

[a] Conditional on $y_{i,t-1}$=1.
[b] Conditional on $y_{i,t-1}$=0.
Notes: Standard errors are shown in parentheses. Estimates (in bold) are obtained from linear probability models of transitions in labor market states on the sample of single childless women and lone mothers. N denotes the number of wave-on-wave state-specific transitions. Other definitions and all time-varying variables used in the estimation are specified in the note to Table 2.

Table 5. The Effect of Welfare Reforms on Other Outcomes – Fixed-Effects Estimates

Outcome	(i) $[\beta_{1999}=\beta_{2000}=\beta_{2001}=\beta\ ]$		(ii)			
	β	β_{1998}	β_{1999}	β_{2000}	β_{2001}	β_{1998}
1. Working 30+ hours per week (N=14,357)[a]	**0.068** (0.020)	**0.034** (0.019)	**0.090** (0.022)	**0.028** (0.024)	**0.051** (0.028)	**0.030** (0.019)
2. Labor force participation (N=14,357)[b]	**0.055** (0.020)	**0.047** (0.019)	**0.057** (0.021)	**0.044** (0.018)	**0.073** (0.027)	**0.050** (0.019)
3. Monthly labor income (N=9,740)[c]	**66.58** (28.25)	**4.62** (9.08)	**62.01** (20.40)	**67.97** (27.36)	**70.22** (31.25)	**4.11** (9.55)
4. FC/WFTC receipt (N=5,283)[d]	**0.113** (0.018)	**0.001** (0.021)	**0.102** (0.019)	**0.117** (0.023)	**0.129** (0.025)	**0.008** (0.021)
5. FC/WFTC monthly award (N=2,337)[e]	**45.65** (13.21)	**11.51** (16.29)	**40.58** (13.35)	**46.72** (17.99)	**49.08** (21.03)	**12.26** (16.49)
6. IS receipt (N=14,357)[a]	**-0.025** (0.012)	**-0.039** (0.015)	**-0.023** (0.012)	**-0.033** (0.019)	**-0.021** (0.022)	**-0.039** (0.016)
7. Paid childcare usage (N=5,283)[f]	**0.028** (0.011)	**0.002** (0.012)	**0.023** (0.009)	**0.029** (0.011)	**0.037** (0.018)	**0.003** (0.012)
8. Weekly childcare costs (N=351)[g]	**16.49** (6.72)	**1.61** (6.92)	**16.62** (7.36)	**15.87** (6.83)	**16.55** (7.51)	**1.59** (7.02)
9. Entry into marriage (N=15,634)[h]	**-0.022** (0.007)	**-0.009** (0.012)	**-0.021** (0.009)	**-0.020** (0.010)	**-0.025** (0.011)	**-0.010** (0.008)
10. Birth rates for lone mothers (N=4,782)[i]	**-0.014** (0.008)	**-0.008** (0.012)	**-0.013** (0.008)	**-0.019** (0.010)	**-0.012** (0.018)	**-0.006** (0.012)
11. Entry into lone motherhood (N= 6,410)[j]	**-0.0018** (0.0015)	**-0.0009** (0.0013)	**-0.0024** (0.0022)	**-0.0012** (0.0017)	**-0.0017** (0.0019)	**-0.0008** (0.0013)

[a] Estimates are obtained from linear probability models with individual fixed effects on the sample of single childless women and lone mothers. Explanatory variables are as in the basic model plus 1998 time dummy in Table 2.

[b] Estimates are obtained from linear probability models with individual fixed effects on the sample of single childless women and lone mothers. The dependent variable takes value one if a woman works (any positive number of hours), and zero otherwise. Explanatory variables are as in the basic model plus 1998 time dummy in Table 2.

[c] Estimates are obtained from linear regression models with individual fixed effects on the sample of single childless women and lone mothers. Monthly earnings are expressed in constant (2001) prices. Explanatory variables included in such models are as in the basic model plus 1998 time dummy in Table 2.

[d] Estimates are obtained from linear probability models with individual fixed effects on the subsample of lone mothers. Explanatory variables included are as in the basic model plus 1998 time dummy in Table 2, except that α_{31} is set to zero.

[e] Estimates are obtained from linear regression models with individual fixed effects on the subsample of lone mothers who report positive monthly amounts of FC/WFTC. The monthly amounts of FC/WFTC are expressed in constant (2001) prices. Explanatory variables included are as in the basic model plus 1998 time dummy in Table 2, except that α_{31} is set to zero.

[f] Estimates are 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. Explanatory variables included are as in the basic model plus 1998 time dummy in Table 2, except that α_{31} is set to zero.

[g] Estimates are obtained from linear regression models with individual fixed effects on the subsample of 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. Explanatory variables included are as in the basic model plus 1998 time dummy in Table 2, except that α_{31} is set to zero.

[h] 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 marries (after which her observations are censored). Multiple entries for the same woman are allowed. Explanatory variables included in such models are as in the basic model plus 1998 time dummy in Table 2.

[i] Estimates are 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 included are as in the basic model plus 1998 time dummy in Table 2, except that α_{31} is set to zero.

[j] Estimates are 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 included are as in the basic model plus 1998 time dummy in Table 2, except that α_{32} is set to zero.

Notes: Standard errors are shown in parentheses. N = number of person-wave observations.

Table 6. The Effects of Welfare Reforms on Other Outcomes by Child's Age Group – Fixed-Effects Estimates

	One child aged:		
Outcome	0-4	5-10	11-18
1. Working 30+ hours per week	**0.149**	**0.052**	**0.068**
	(0.030)	(0.023)	(0.021)
2. Labor force participation	**0.112**	**0.055**	**0.074**
	(0.030)	(0.026)	(0.024)
3. Monthly labor income	**126.37**	**97.24**	**74.83**
	(45.81)	(38.30)	(37.94)
4. FC/WFTC receipt	**0.176**	**0.163**	**0.104**
	(0.034)	(0.026)	(0.024)
5. FC/WFTC monthly award	**54.38**	**48.40**	**39.89**
	(19.15)	(20.59)	(16.68)
6. IS receipt	**-0.031**	**-0.025**	**-0.020**
	(0.014)	(0.013)	(0.022)
7. Paid childcare usage	**0.043**	**0.030**	
	(0.014)	(0.012)	
8. Weekly childcare costs	**15.16**	**13.56**	
	(5.97)	(5.38)	
9. Entry into marriage	**-0.035**	**-0.005**	**-0.008**
	(0.013)	(0.021)	(0.011)
10. Birth rates for lone mothers	**-0.021**	**-0.003**	**0.008**
	(0.010)	(0.015)	(0.012)

Notes: All estimates reported in this table are obtained from the equation (8) type of regressions with the same samples and models as those used to obtain the estimates under specification (i) in Table 5. For details see the notes to that table. Standard errors are shown in parentheses.

Table 7. The Effect on Lone Mothers' Labor Supply of the Change in Child Rate Components of the FC/WFTC between 1998 and 1999

Outcome and sample	(i) Child aged 10 or less in 1998	(ii) Child aged 11 or more in 1998	(iii) Child aged 11 or less in 1999	(iv) Child aged 12 or more in 1999	(v) Wald estimates (std. error) [(iii)-(i)]-[(iv)-(ii)] (no controls)	(vi) Difference (std. error) [(iii)-(i)]-[(iv)-(ii)] (with controls)
Working 16+ hours per week						
Sample 1	0.429	0.421	0.534	0.512	0.014 (0.076)	0.0004 (0.076)
Sample 2	0.459	0.434	0.548	0.528	-0.005 (0.054)	-0.015 (0.055)
Working 30+ hours per week						
Sample 1	0.317	0.303	0.353	0.332	0.007 (0.075)	0.004 (0.075)
Sample 2	0.284	0.326	0.360	0.394	0.008 (0.050)	-0.003 (0.049)

Notes: Sample 1 is the subsample of lone mothers who have children aged 10 or 11 in 1998 and who are observed both in 1998 and 1999 (there are 50 women with children aged 10 in 1998, and 57 with children aged 11 in 1998). Sample 2 is the subsample of lone mothers who have children aged 6 to 14 in 1998 and who are observed both in 1998 and 1999 (there are 322 women with children aged 6-10 in 1998, and 244 women with children aged 11-14 in 1998). The total number of person-wave observations over which the estimates in both columns (v) and (vi) have been computed is 214 and 1,132 for Sample 1 and Sample 2, respectively. The differences in column (vi) are obtained from regressions that include controls for age, education, race, housing tenure, region of residence, and a child's age dummy variable.

Table 8. The Effect of Welfare Reforms on Lone Mothers' Transition Probabilities of Working 16 or More Hours per Week and Using Paid Childcare

Type of transition and model	(i) $[\beta^{(j)}_{1999} = \beta^{(j)}_{2000} = \beta^{(j)}_{2001} = \beta^{(j)}]$		(ii)			
	β	β_{1998}	β_{1999}	β_{2000}	β_{2001}	β_{1998}
A. From nonworking to working 16+ and using paid childcare[a] (N=2,065)						
a. No distinction by child's age group						
	0.032	**-0.001**	**0.029**	**0.027**	**0.038**	**0.005**
	(0.015)	(0.010)	(0.014)	(0.015)	(0.019)	(0.010)
b. Distinguishing by child's age group						
One child aged 0-4	**0.037**	**0.003**	**0.046**	**0.027**	**0.037**	**0.006**
(j=1)	(0.017)	(0.008)	(0.020)	(0.018)	(0.018)	(0.009)
One child aged 5-10	**0.031**	**0.005**	**0.036**	**0.023**	**0.034**	**0.008**
(j=2)	(0.014)	(0.010)	(0.015)	(0.018)	(0.016)	(0.012)
One child aged 11-18	**0.003**	**-0.012**	**0.007**	**0.004**	**-0.001**	**0.002**
(j=3)	(0.022)	(0.018)	(0.020)	(0.019)	(0.021)	(0.021)
B. From working 16+ without paid childcare to working 16+ and using paid childcare (N=1,126)						
a. No distinction by child's age group						
	0.036	**-0.002**	**0.038**	**0.031**	**0.035**	**-0.006**
	(0.014)	(0.021)	(0.016)	(0.017)	(0.016)	(0.021)
b. Distinguishing by child's age group						
One child aged 0-4	**0.045**	**0.004**	**0.064**	**0.041**	**0.042**	**0.005**
(j=1)	(0.019)	(0.018)	(0.027)	(0.020)	(0.019)	(0.019)
One child aged 5-10	**0.034**	**-0.003**	**0.039**	**0.032**	**0.031**	**-0.005**
(j=2)	(0.016)	(0.020)	(0.020)	(0.018)	(0.017)	(0.021)
One child aged 11-18	**0.026**	**-0.005**	**0.020**	**0.017**	**0.028**	**-0.008**
(j=3)	(0.018)	(0.023)	(0.021)	(0.023)	(0.022)	(0.021)

[a] The nonworking (origin) state includes women who work less than 16 hours per week and (if working) do not use paid childcare.

Notes: Robust standard errors are shown in parentheses. Estimates (in bold) are obtained from linear probability models of transitions in labor market states on the subsample of lone mothers. N denotes the number of wave-on-wave state-specific transitions. Other definitions and variables used in the estimation are specified in note f to Table 5.

Table 9. The Effect of Changing Child's Age between 1998 and 1999: Lone Mothers' Labor Supply

Outcome	(i) Child aged 3 in 1998 (N=42)	(ii) Child aged 5 in 1998 (N=42)	(iii) Child aged 4 in 1999 (N=42)	(iv) Child aged 6 in 1999 (N=42)	(v) Wald estimates (std. error) [(iii)-(i)]-[(iv)-(ii)] (no controls) (N=168)	(vi) Difference (std. error) [(iii)-(i)]-[(iv)-(ii)] (with controls) (N=168)
Working 16+ hours per week	0.238	0.333	0.429	0.476	0.048 (0.019)	0.029 (0.016)
Working 30+ hours per week	0.098	0.166	0.293	0.310	0.051 (0.019)	0.037 (0.017)

Notes: The differences in column (vi) are obtained from regressions that include controls for age, education, race, housing tenure, region of residence, and a child's age dummy variable. N is the number of lone mothers in columns (i)-(iv) and the number of person-wave observations in columns (v) and (vi).

Table 10. The Effect of Welfare Reforms on Lone Mothers' Labor Supply by Child's Specific Age (N=10,584)

	Working 16+ hours per week			Working 30+ hours per week		
	(i)	(ii)		(i)	(ii)	
Child' age	β	β	β_{1998}	β	β	β_{1998}
Age 2	**0.025**	**-0.003**	**-0.017**	**-0.001**	**0.010**	**-0.038**
	(0.097)	(0.085)	(0.083)	(0.079)	(0.081)	(0.089)
Age 3	**0.152**	**0.156**	**0.044**	**0.114**	**0.121**	**0.053**
	(0.099)	(0.100)	(0.076)	(0.102)	(0.101)	(0.078)
Age 4	**0.182**	**0.191**	**0.015**	**0.183**	**0.187**	**0.031**
	(0.075)	(0.076)	(0.091)	(0.076)	(0.077)	(0.092)
Age 5	**0.022**	**0.032**	**0.077**	**0.009**	**0.008**	**-0.019**
	(0.054)	(0.053)	(0.071)	(0.055)	(0.057)	(0.072)
Age 6	**0.035**	**0.037**	**-0.009**	**0.026**	**0.025**	**0.023**
	(0.050)	(0.059)	(0.088)	(0.051)	(0.054)	(0.090)
Age 7	**0.118**	**0.130**	**-0.010**	**0.082**	**0.101**	**0.020**
	(0.049)	(0.052)	(0.080)	(0.050)	(0.053)	(0.081)
Age 8	**0.062**	**0.059**	**0.024**	**0.027**	**0.025**	**0.017**
	(0.081)	(0.077)	(0.070)	(0.055)	(0.059)	(0.078)
Test of equality (p-value) of:[a]						
$\beta^{(4)} = \beta^{(5)}$	0.027	0.031		0.019	0.023	
$\beta^{(4)} = \beta^{(6)}$	0.034	0.036		0.038	0.037	

Notes: Robust standard errors are shown in parentheses. Estimates (in bold) are obtained from linear probability models on the sample of single childless women and lone mothers. Other conditioning variables are: a quartic polynomial in age, one linear time trend for lone mothers, and one linear time trend for single childless women; and 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), lone mother status, child's age group, and interactions between mother's age and the educational group dummies, mother's age child's age group, and mother's education and child's age group. N = number of person-wave observations.

[a] These p-values come from F-tests of equality between the relevant coefficients at different ages of children. Under specification (i), the tests have 1 and 10538 degrees of freedom, while under specification (ii) the tests have 1 and 10531 degrees in freedom.

Table A1. 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
Monthly labor income conditional on working positive hours (in 2001 pounds)	1,036 (1,013)	608 (704)
FC/WFTC receipt (all lone mothers)		0.236
FC/WFTC receipt (all working lone mothers)		0.482
FC/WFTC monthly award conditional on reporting positive values (in 2001 pounds)		168.45 (148.90)
IS receipt	0.078	0.386
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		0.085
Birth rates for lone mothers		0.037
Entry into lone motherhood	0.013	
Main explanatory variables		
Age	31.319 (12.410)	28.541 (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.