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Does more free childcare help parents work more?☆

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ABSTRACT

Many governments are considering expanding childcare subsidies to increase the labour force participation of parents (especially mothers) with young children. In this paper, we study the potential impact of such a policy by comparing the effects of offering free part-time childcare and of expanding this offer to the whole school day in the context of England. We use two different strategies exploiting free childcare eligibility rules based on date of birth. Both strategies suggest that free part-time childcare only marginally affects the labour force participation of mothers whose youngest child is eligible, but expanding from part-time to full-time free childcare leads to significant increases in labour force participation and employment of these mothers. These effects emerge immediately and grow over the months following entitlement. We find no evidence that parents adjust their labour supply in anticipation of their children's entitlement to free childcare.

1. Introduction

Over the last two decades, most OECD countries have introduced policies that make childcare cheaper or more readily available, with the aim of increasing parental labour supply and/or promoting child development. Despite these efforts, the cost of childcare is still a big concern for many parents, potentially hindering their labour market attachment. In recent years, these concerns have led several countries to expand the generosity of their childcare subsidies, e.g. by extending childcare subsidies to younger children or by increasing the number of hours of subsidised care available.[1] But, in other countries, how much childcare should be subsidised remains an important policy question. In the US, for example, this issue was highly debated in the 2020 presidential election, with several Democratic candidates proposing major plans to expand child care subsidies for families with young children.

The existing empirical literature offers a wide range of estimates of the impact of part-time and full-time childcare subsidies on maternal labour supply (see Cascio et al. (2015) and Cattan (2016) for reviews). However, with the exception of a few studies we discuss later on, most focus on estimating the impact of offering either subsidised part-time childcare or subsidised full-time childcare compared with offering nothing. As such, this literature is limited in its ability to inform the likely impact of extending the offer of free or subsidised childcare to cover more hours of the day. To do so would require comparisons to be made across countries or time periods with very different contexts. Moreover, parents affected by extensions of childcare subsidies likely differ from

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[1] For example, in 2002, Sweden passed a major childcare price reform, which lowered further an already highly-subsidized price of childcare, and Norway followed with a similar reform (Lundin et al., 2008). In England, the offer of free childcare for 3 and 4 year olds was expanded from 15 to 30 h per week for working families in September 2017, a change we do not exploit in this paper (because too little time has elapsed to detect any impact precisely). In 2018, Berlin abolished childcare place fees for children, even under the age of 1, and Lower Saxony and Hesse followed suit in 2019. Thuringia, in contrast, effectively raised childcare prices by subsidizing not using childcare (Gathmann and Sass, 2018).

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those affected by their introduction, and subsidies are likely to have non-linear effects on parental employment, for example because of inflexible job contracts. Further, amongst parents who are already in work, extending the subsidy would have a priori ambiguous effects on the number of hours worked, as its impact would depend on the relative strengths of the income and substitution effects when trading-off between work and leisure. Thus, even in contexts where the introduction of childcare subsidies did encourage some parents to work or work longer hours, it is not clear that extending them further would yield any further increase in labour supply.

The main contribution of this paper is to shed light on this issue by evaluating the impact on mothers' and fathers' labour supply of initially offering pre-school children in England free, half-day childcare and then increasing this offer to the whole of the school day when they start formal schooling. We make this comparison based on results obtained using the same datasets and within the same institutional setting and time-period. The paper is amongst the first to consider the effects of such an expansion for pre-school aged children, a margin of particular relevance to policy-makers interested in increasing labour force participation of mothers with young children. A distinct feature of our analytical approach is that we consider how parents' labour supply responses to the provision of free childcare evolve with the duration of the subsidy and the extent to which anticipation effects might be responsible for the patterns we see.

Eligibility for free childcare – including in England – usually depends on the child's age. As such, the main identification challenge is to separately identify the effect of eligibility for free childcare from the independent effect of child's age on parental labour supply. To overcome this challenge, we exploit birth date-based rules governing children's entitlement to free part-time and full-time childcare. Specifically, in England, children are eligible for a free part-time childcare place at the start of the school term after they turn three (in either September, January or April), and most children are eligible to start full-time school in the September after they turn four (we refer to full-time school as full-time childcare for the rest of the paper). These rules mean that children gain entitlement to free care at different ages and remain entitled for differing amounts of time, thus generating plausibly exogenous variation in eligibility for free childcare and duration of entitlement conditional on age.

We exploit these rules to implement two empirical strategies. First, we follow a number of other papers in this literature in adopting a Regression Discontinuity (RD) design (e.g. Berlinski et al., 2011; Fitzpatrick, 2010; Goux and Maurin, 2010). In our case, the impact of eligibility for free part-time or full-time childcare is identified by comparing the outcomes of parents whose children become eligible for a particular type of free care at a given point in time with those of parents whose children become eligible a term (in the case of part-time care) or a year (in the case of full-time care) later, simply because they are born a few days later. Following Gelbach (2002) and Fitzpatrick (2010), we implement this approach using Census data – specifically data from the 2011 UK Census. Like these US studies, because the UK Census date falls in late March, this enables us to estimate the impact of free full-time childcare relative to free part-time childcare on parental labour supply some seven months after children first become entitled to free full-time care. Because children become entitled to free part-time childcare each term (roughly every four months) rather than each year, however, the same data enables us to investigate whether the impact of entitlement to free part-time childcare varies by duration of exposure, as we are able to compare the parents of children who have been entitled to free part-time childcare for zero vs. one, one vs. two, two vs. three, and three vs. four terms at the time of the Census.

By comparing the outcomes of individuals whose children are born very close to the cut-off dates, and hence unlikely to differ in unobserved ways, the RD approach provides a clean way to identify the causal impact of entitlement to free childcare. However, as outlined above, it only enables us to assess how the effects of entitlement vary with duration of exposure to free part-time care, not free full-time care. Moreover, a potential limitation of our – and indeed all – RD approaches in this literature is that the estimates are specific to parents of children born at particular times of the year and may therefore not reflect average effects. This could be the case if, as emphasised by Buckles and Hungerman (2013) and Clarke et al. (2019), mothers trying to conceive at different times of the year differ in observed and unobserved ways, such as family background or preferences regarding family and work.

To address these concerns, we supplement the RD analysis with a second panel data approach. We implement it using the UK Labour Force Survey (LFS), which collects labour supply information on a nationally representative sample of households every quarter, for up to 5 quarters.[2] These frequently repeated observations enable us to identify the treatment effects for children born in all months of the year from within-parent changes in labour supply as their children's entitlement to free childcare changes over time. This allows us to consider heterogeneity in the impact of entitlement to free full-time as well as free part-time childcare by duration of exposure and to estimate average effects across children born in all months of the year. The LFS sample size is too small to identify separate effects at all relevant RD cut-offs, so we broaden the windows around the birth date-based discontinuities in entitlement used in the RD strategy above to include children born throughout the year. As the parents of these children are more likely to differ from each other than parents of children born just before and just after particular cut-offs, we include parent-level fixed effects to control for time-invariant differences between them. To our knowledge, such an approach has not been used in the context of evaluating the impact of childcare policies on parental labour supply, although Black et al. (2011) combine birthday-based rules governing entitlement to start school and family fixed effects to estimate the impact of school starting age on children's IQ in Norway.

Our main findings can be summarised in four points. First, offering free childcare never affects the labour market outcomes of fathers and only affects the labour market outcomes of mothers who have no younger child. Second, the provision of free part-time childcare has at most a small effect on the labour force participation of those mothers. Third, offering free childcare to cover a full school day instead of a half day significantly increases their labour force participation and employment. Our estimates suggest that mothers are at least 3 percentage points (ppts) more likely to be in the labour force and 1 ppt more likely to be in paid work in the first term after their youngest child is offered free full-time childcare instead of free part-time childcare. Fourth, mothers' labour supply response to childcare subsidies varies by duration of exposure – the labour force participation impact of free full-time childcare is almost twice as large by the end of the first year of full-time entitlement as it is in the first term. The employment impact is more than three times as large, corroborating the hypothesis that it takes time for mothers to enter the labour force and find a paid job.

To better understand these results, we investigate how eligibility for free part-time and full-time childcare affects the take-up of formal and informal childcare by drawing on another dataset, the Family Resources Survey, with rich childcare information. We find that the entitlement to free part-time childcare increases the use of formal, subsidisable care. However, it crowds out the use of informal childcare, so that there is little change in the total amount of time that children spend in any form of childcare. In contrast, the rise in the use of subsidisable childcare following entitlement to free full-time childcare does not entirely crowd out the use of other forms of childcare. These results are fully consistent with the small labour supply response to part-time eligibility and the stronger response to full-time eligibility we estimate in both datasets.

[2] The Labour Force Survey is most similar in design to the Current Population Survey (CPS) in the US, with the exception that it follows households quarterly instead of monthly. We were also able to obtain access to children's exact date of birth.

Our paper makes several contributions to the literature on the impact of childcare policies on parental labour supply. First, it offers evidence of the impact on both mothers' and fathers' labour supply of increasing the provision of free childcare from half-day to full-day care amongst children under five. This contrasts with the vast majority of existing studies on this topic, which focus on mothers only[3] and either study the impact of offering subsidised or free childcare compared to offering nothing or else consider the impact of extending childcare subsidies for older children. Some of the very few that do consider the impact of lengthening the number of hours of care provided include (Berthelon et al., 2015) and Shure (2019) who evaluate the impact of policies to increase the length of the primary school day in Chile from about 5.5 to 7.5 h and in Germany from about 5 to 7 h, respectively. Both papers find positive effects of the reform on mothers' labour force participation. However, the results are not directly comparable with our setting as the children affected by the reforms are older (6–13 and 6–10 respectively, compared to age 4–5 in this paper) and the extension offered is lower at around 2 h per day, compared to around 3.5 h in this paper. Similarly, our findings do not necessarily predict labour supply responses of eligibility at earlier ages.[4] More similar to our setting is that studied by Dhuey et al., 2019 and Dhuey et al. (2020) who exploit reforms that lengthened the kindergarten school day (affecting children aged 4–5) in Ontario's public-funded schools in the late 1990s in French-speaking schools, and then from 2010 in English-speaking schools from about 2.5 to about 6.6 h per day. Dhuey et al. (2020) find the late 1990s change led to a large rise in employment amongst French-speaking single mothers, and (Dhuey et al., 2019) found that the more recent reform had no impact at the extensive margin, but did increase weekly average hours worked by just under 2.[5]

The second contribution of our paper is to investigate how the impact of childcare subsidies varies by duration of exposure. In doing so, we add to a small set of papers interested in how mothers' labour market behaviour following receipt of a childcare subsidy evolves over time (Lefebvre et al., 2009; Nollenberger and Rodriguez-Planas, 2015) and we show that this matters for our understanding of the effect of these policies. This is important because most existing studies have estimated the impact of childcare subsidies on maternal labour market outcomes at a single point in time following the child's eligibility (typical amongst regression discontinuity approaches, such as Goux and Maurin (2012) and Fitzpatrick (2010)) or its average impact across several months or years of eligibility (more common amongst studies that exploit staggered expansion of childcare provision, such as Havnes and Mogstad (2011) and Berlinski and Galiani (2007)).

Finally, we develop a panel data-based identification strategy to complement the more traditional RD approach commonly used to identify the impact of free childcare on parental labour supply using birthdate-based eligibility rules. Its main advantage compared to the RD approach is that it allows us to recover average impacts of free part-time and full-time childcare in every school term following the receipt of the subsidy across children born in every month of the year, albeit at the cost of an arguably stronger identifying assumption, namely that the outcomes of parents of children born across the year – as opposed to just at the discontinuity – do not differ in time-varying ways for which we cannot control. The fact that our results are broadly consistent across the two approaches makes our findings particularly robust and provides confidence in the validity of the panel data-based approach.

Moreover, by implementing the panel data-based approach using the LFS – which includes a rich set of covariates and covers all years between the late 1990s and early 2010s – we are also able to conduct heterogeneity analysis and robustness checks in ways the Census data would not allow us to do. Among others, we test for the possibility that parents make labour supply decisions in anticipation of receipt of the subsidy, which could under- or over-estimate the true impacts of being entitled to some free childcare compared to being entitled to nothing, depending on how parents respond.[6] While this potential issue is common to all designs based on known cut-off rules in the related literature, its implications have been under-explored to date.[7] We present empirical evidence suggesting that, in our case, anticipation effects of at least one year before eligibility are not a concern for the interpretation of our results.

The remainder of this paper is organised as follows. Section 2 provides background on childcare policy in England. Section 3 reflects on the policy effects we might expect to find and Section 4 the describes our empirical strategies and data. Section 5 presents our RD results, while Section 6 presents our panel data results. Section 7 discusses our results in relation to the literature and presents additional analysis of the effect of the policies considered on childcare use. Section 8 concludes.

2. Institutional background

2.1. Free part-time childcare for 3- and 4-year-olds

In the mid to late 1990s the UK had a relatively low maternal employment rate: only 57% of mothers of children aged 0–6 were in work, and this proportion was lower for lone mothers (40% in work) and low educated mothers (44% in work).[8] Together with the perception that childcare was not affordable for many families, this has contributed to a substantial increase in public support for pre-school childcare in England (and the rest of the UK) over the past 20 to 25 years.

Although there are other forms of childcare support on offer, in England the largest proportion of funding goes to the "free entitlement" policy, which we exploit in this paper.[9] As part of this policy, since the early 2000s, all three and four year olds in England have been entitled to

[3] Examples of papers that investigate impacts of childcare on fathers' employment are Felfe et al. (2016) and Andresen and Havnes (2019).

[4] A related paper by Felfe et al. (2016) examines the impact of providing after school care for 4–12 year olds in Switzerland and finds impacts on mothers' full-time work but not employment overall.

[5] Related to this paper, Lundin et al. (2008) study a policy change that introduced a price cap on already highly subsidised childcare for children aged 1–9, thereby halving the average hourly rate from 14.7 SEK (USD1.75 or GBP 1.22 at today's rates) and show that these changes led to increased attendance mostly among children of unemployed parents and parents on parental leave, and no impact on mothers' employment and hours of work (amongst those who are working). We also note that Cannon et al. (2006) estimate the impact of attending full-day kindergarten versus half-day kindergarten on maternal work using data from the Early Child Longitudinal Study-Kindergarten Class of 1998–1999. To address parental selection into full-day versus half-day kindergarten, they use state (but not time) variation in policies on full-day kindergarten programs as an instrument for the likelihood that a student will attend a full-day program. However, they warn that their results should be viewed with caution given that they find only mixed evidence suggesting the validity of their instruments.

[6] For example, the difficulties in finding good part-time jobs means that parents might move into full-time work when they become entitled to part-time childcare, in the knowledge that any childcare that they buy will soon become free. On the other hand, the same difficulties might mean that parents will not look for work until they become entitled to free full-time childcare.

[7] Anticipation effects are not an issue in paperes which rely on policy changes that were not expected (Baker et al., 2008; Bauernschuster and Schlotter, 2015) or where admission to childcare is uncertain (Drange and Havnes, 2019). Where they are an issue, anticipation effects in the context of the link between entitlement to childcare subsidies and parental labour have not received a lot of empirical attention but have been explored in the context of other entitlements (see Berg et al. (2020) for a recent example).

[8] Source: author's calculations based on the Quarterly Labour Force Survey for 1992 to 2000. Low educated is defined as those with less than A-levels, a group that is the equivalent of those without a high school degree in the US.

[9] Other forms of childcare support on offer during the 2000s include a refundable tax credit that subsidises up to 80 percent of spending on formal childcare amongst low- to middle-income working families (available throughout the UK), as well as a scheme to allow employers to pay childcare vouchers that are free of personal income tax and social insurance contributions (also available through-

receive free part-time childcare before entering full-time primary education (which they would typically do between the age of 4 and 5, as we discuss later).[10] Crucial to our identification of policy impacts are the various discontinuities in eligibility caused by date-of-birth admission rules. Children become eligible for a free part-time childcare place at the start of the academic term after they turn three (well after statutory maternity leave ends when the child turns one). This means that children born between 1 January and 31 March ('spring-borns') are eligible for a free place from 1 April of the year they turn three; children born between 1 April and 31 August ('summer-borns') are eligible for a free place from 1 September of the year they turn three; and children born between 1 September and 31 December ('autumn-borns') are eligible from 1 January of the calendar year in which they turn four. Children remain entitled to free part-time childcare into their fourth year of life until they enter full-time primary education, the policy we exploit in this paper to identify the impact of extending care from part-time to full-time hours.

Parents can use their entitlement either in one of a limited number of state-run childcare settings or in a childcare facility run by the private sector.[11] Eligibility rules are the same across both sectors. By 2013, the end of the period we analyse, 93% of children used at least some of the hours to which they are entitled, and the majority of these children used all of the hours to which they are entitled (Department for Education, 2013). Only in private nurseries can parents pay for additional hours on top of their entitlement. Indeed, a marked difference between England and many other countries is the existence of a private market for childcare, with 60% of families with a two year old already paying for some form of private childcare before their child is entitled to free care (see Appendix Table 1).[12] This means that the free entitlement can effectively be viewed as a price subsidy, rather than as a policy that hugely increased the availability of childcare places, as is often studied in other countries.

While children are legally entitled to a free part-time place at the start of the term after they turn three – and have been since the early 2000s - there are two ways in which capacity constraints might potentially weaken the effect of the entitlement on parental labour supply in our analysis. First, children born in different terms of the year may face differential chances of securing a place at nursery. This is because nursery places in England tend to become available from September, the month in which most children start full-time schooling and therefore vacate places in nurseries. This is also the month in which summer-born children first become entitled to free part-time childcare. This could imply that the parents of autumn- and spring-born children may not be able to secure a place at their preferred nursery as soon as their children become entitled. Second, although places should have been universally available from 2004, full coverage of funded places was not achieved until about 2007 (see Blanden et al. (2016) who exploit this feature to identify effects of childcare availability on child outcomes). In the presence of such capacity constraints, we would expect to underestimate the impact of childcare eligibility. In Section 6.3 we present two robustness checks that assess whether these types of capacity constraints affect our results. Overall, we find no evidence that it is the case.

2.2. Free full-time childcare for 4-year-olds

Parents in England are statutorily obliged to send their child to school from the school term that begins after the child's fifth birthday (the 'statutory school age'), earlier than in most OECD countries. However, schools have the discretion to admit children earlier than this, and almost all children in England are able to attend full-time school (covering about 6.5 h a day, or 30 to 35 h a week, depending on school policy, for 39 weeks a year) before the statutory school age. Indeed, in 2012 more than 99% of children in England started school in an area which allowed them to do so in the September after they turned four, up from around 80% in the early 2000s.[13] Parents do not have to send their child to school earlier than the statutory school age, but the vast majority of children do start school in the September after they turn four.[14]

This policy introduces further variation in entitlement to childcare which is crucial to our identification strategies. The fact that most children start school in the September after they turn four generates variation across those born in different months of the year in both the age at which children become entitled to full-time care and the number of terms of part-time care that they can receive. For example, children born one day apart on 31 August and 1 September 2011 would be eligible for a free part-time nursery place four months apart (1 September 2014 vs 1 January 2015), and a free full-time school place 12 months apart (1 September 2015 vs 1 September 2016). This also means that children born on 31 August are only eligible to receive three terms of part-time care before starting full-time education, whereas children born on 1 September are eligible for five terms of free part-time care before starting school. Spring-borns are eligible for four terms of free part-time childcare.

Fig. 1 illustrates the variation in access to free part-time and full-time childcare created by the different eligibility rules for children born in each month of the year. It shows the ages at which these children become eligible for their first, second, third, and for some children fourth and fifth, terms of part-time childcare, and the ages at which they become eligible for different terms of full-time care.[15] Although there is no age at which we observe children with all possible entitlements to free childcare, it is the case that, for every possible age in months from 36 to 60, we observe children in between two and four different possible entitlement statuses. As we will elaborate, we use this variation to

out the UK). See Brewer et al. (2014) for a more detailed analysis of the childcare policy landscape in England.

[10] This entitlement has been in place for all four-year-olds since 2000 and for all three-year-olds since 2004. When the policy was first introduced, it offered 2.5 h of free childcare per day (12.5 h per week) for 33 weeks a year. This entitlement was extended to 38 weeks a year in 2006 and to 15 h a week in 2010. Since 2010, it can also be taken with greater flexibility: in some settings, families can now use the hours across a minimum of three days, making it easier to combine with work.

[11] The existence of these state-run institutions providing childcare pre-dates the policy we study: since the early 1990s, some local authorities in England have been providing free pre-school education in nursery classes in schools or in stand-alone nursery schools, and these use the same date-of-birth admission rules as the ones we exploit in this paper. Because the variation we exploit in this paper is by age and term of birth rather than by policy period, the existence of state-run institutions does not affect the interpretation of our results. We do however focus on the period from 2004 when estimating the impact of eligibility for free part-time childcare because places for 3-year-olds were only universally available from that year.

[12] As we describe later in the paper, we use the Family Resources Survey, a nationally representative cross-sectional sample of households between 2005 and 2013, to describe patterns of childcare use by age of the youngest child and explore how childcare use changes as children get entitled to free part-time and full-time childcare. Before the age of 3, formal childcare is almost entirely provided in the private sector.

[13] Source: authors' calculations using administrative data on children attending state schools in England from the National Pupil Database. Schools which do not offer all children the opportunity to start school in the September after they turn four instead operate dual or triple entry point systems, with date-of-birth cut-offs determining which children start in which term. Our results are robust to accounting for the most common school admissions policy in operation in the local area.

[14] One reason for this is that caps on class sizes mean that parents often cannot secure their child's place at a particular school if they defer entry.

[15] The first, second, etc. terms are not always the same duration for children born in different months of the year because children both in different months start their entitlement in different academic terms, and the three academic terms have different length (the Autumn term is 4 months long, the Winter term is 3 months long, and the Summer term is 5 months long).

Table 1
Descriptive statistics of the initial and final LFS samples.

	(1)			(2)		
	Initial sample			Final sample after sampling decisions		
Sample of mothers	Mean	Std. dev.	N	Mean	Std. dev.	N
In labour force	0.610		296,866	0.614		276,018
In work	0.568		296,866	0.572		276,018
Works 1–15 h/wk	0.107		294,536	0.109		273,920
Works 16–29 h/wk	0.237		294,536	0.240		273,920
Works 30+ h/wk	0.220		294,536	0.218		273,920
Usual weekly hours	14.305	(15.281)	294,536	14.319	(15.212)	273,920
Looking for work	0.049		296,866	0.049		276,018
Age	33.064	(6.070)	294,830	33.107	(6.037)	276,018
Has a partner	0.773		296,866	0.777		276,018
Non white	0.150		296,341	0.144		275,984
Low education (< A-levels)	0.507		296,262	0.502		275,703
Number of kids under 19	1.976	(0.992)	296,866	1.978	(0.988)	276,018
Age of youngest child	2.193	(1.676)	295,617	2.203	(1.655)	276,018
Sample of fathers						
In labour force	0.951		229,498	0.953		213,637
In work	0.909		229,498	0.912		213,637
Works 1–15 h/wk	0.007		224,803	0.007		209,433
Works 16–29 h/wk	0.038		224,803	0.037		209,433
Works 30+ h/wk	0.862		224,803	0.866		209,433
Usual weekly hours	38.403	(15.597)	224,803	38.525	(15.451)	209,433
Looking for work	0.041		229,498	0.040		213,637
Age	36.447	(6.499)	228,052	36.488	(6.460)	213,637
Has a partner	1.000		229,498	1.000		213,637
Non white	0.146		229,099			213,613
Low education (< A-levels)	0.407		228,098	0.402		212,657
Number of kids under 19	1.971	(0.966)	229,498	1.973	(0.960)	213,637
Age of youngest child	2.112	(1.668)	228,611	2.122	(1.647)	213,637

Note: Sample consists of mothers and fathers observed with a child aged 0–6 between January 2000 and December 2013 in the Labour Force Survey.

estimate the impact of entitlement to different types of free care on the labour supply of their parents.

3. Likely effects of the policy

Demand for non-parental childcare is driven by a variety of factors. It could be that childcare raises utility directly, by increasing leisure time, or indirectly, because it is beneficial for children. It might also be a derived demand: someone has to look after the children while parents are working. Parents' labour supply decisions are thus likely to be affected not only by the wages on offer, their own preferences for working, and the labour supply decisions of their partner (if they have one), but also the different types of available care, as well as the price and quality of that care, and other factors related to the convenience of care. Depending on their location and family circumstances, parents may be able to choose between a number of non-parental childcare options, including informal care provided by family or friends and a range of different types of formal care, which must typically be paid for.

Different interventions in the childcare market will affect parental demand for (different types of) childcare in different ways. For example, supply-side interventions, such as building new public childcare facilities or regulating staff-child ratios may affect quality and/or accessibility of formal childcare, with demand potentially changing in response, depending on parents' preferences. Many, perhaps the majority, of interventions, however including those studied in this paper affect the price of formal childcare (see, for example, Blau and Hagy, 1998; Mumford et al., 2020; Powell, 2002, for theoretical discussions of models jointly estimating labour supply and childcare decisions in response to different types of childcare subsidies).

In our case, the price of the first 15 h of formal childcare is effectively reduced to zero at the beginning of the term after a child turns three, with the price of the next (roughly) 15 h of care then falling to zero in the September after the child turns four. The price of formal childcare is assumed to be unchanged for all hours above the relevant cut-off. To put this in context, the average hourly rate charged by formal childcare providers in England for care for 3 and 4 year olds was roughly £5 in 2019.[16] If parents would otherwise have paid for 30 h of care per week for 52 weeks a year, the part-time subsidy (of 15 h per week, 38 weeks a year) amounts to around £55 per week (just over a third of total annual childcare costs), and the full-time subsidy to around £110 per week (over two thirds of total annual childcare costs). This compares to median weekly earnings of between £500 and £600 for women aged 25–39 (and between around 550 and 650 for men of the same age).[17] The size of the subsidy is thus large enough to make a difference to parental labour supply decisions at the margin, but is highly unlikely to affect other margins of response, such as fertility or partnership status.

How will parents respond to each of these changes, and in particular how will their use of childcare and their labour supply be affected? In a static setting such as if these policies were introduced unexpectedly we might expect parents not currently using any form of childcare to start using formal childcare, assuming they are not constrained and do not have strong preferences against doing so. For these (non-working) parents, the entitlement might open up the possibility of moving into work if they are able to find a job that fits within the free hours, or if their net earnings (after paying for any childcare beyond the free entitlement) rises above their reservation earnings.[18]

[16] Source: https://assets.publishing.service.gov.uk/government/uploads/system/uploads/attachment_data/file/845081/SCEYP_2019_LA_Fees_Report_Nov19.pdf.

[17] Source: https://www.ons.gov.uk/employmentandlabourmarket/peopleinwork/earningsandworkinghours/bulletins/annualsurveyofhoursandearnings/2019.

[18] As the entitlement only provides free childcare during term-time, parents would have to pay for additional hours in weeks outside term-time, as well as any additional hours per week.

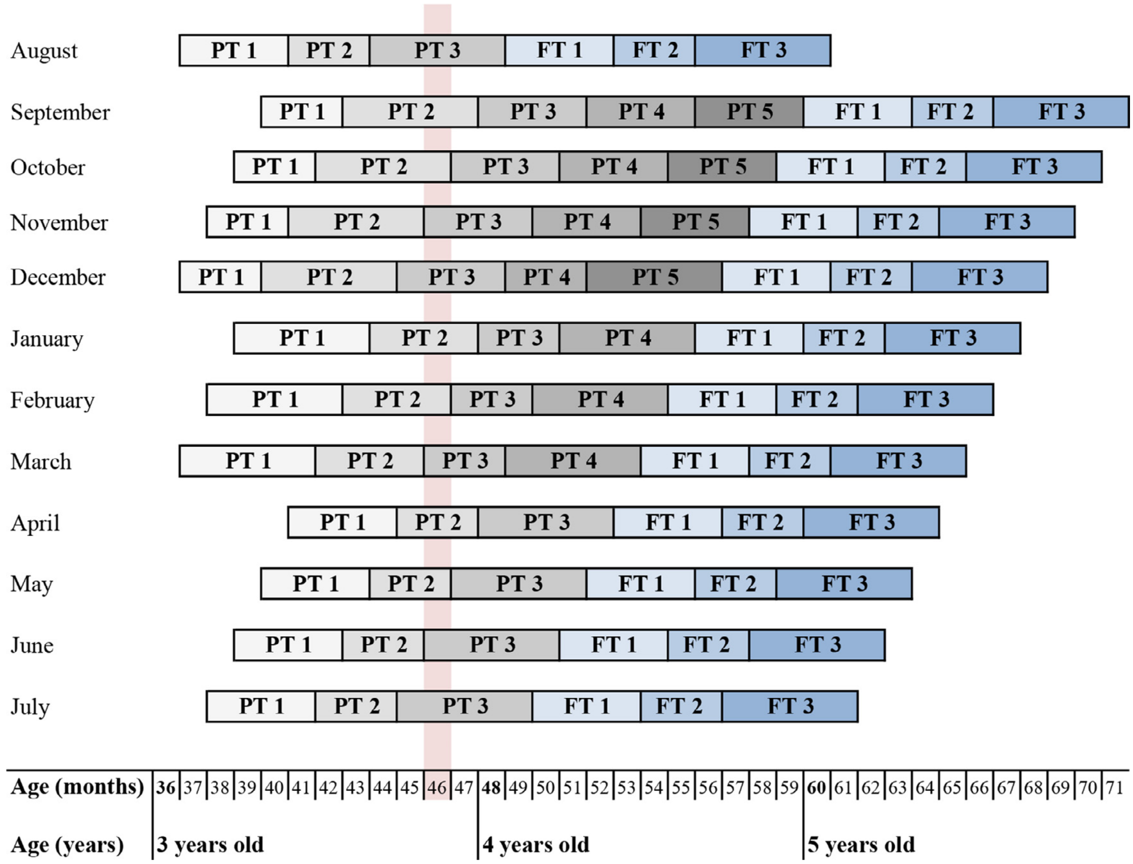

Fig. 1. Terms of entitlement to free part-time (PT) and full-time (FT) childcare for children born in different months of the year. *Notes:* This figure shows the age (in months) children born in different months are when they are in different terms of entitlement to free part-time (PT) and full-time (FT) childcare. PT1 refers to the first school term of entitlement to free part-time childcare, PT2 to the second school term, etc. The red vertical line exemplifies that children born in different months are in different terms of entitlement to PT childcare at the same age (46 months). (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

For those currently using between 0 and 15 (or 15 and 30) hours of formal childcare per week (which they would previously have paid for), the income effect would suggest a reduction in childcare use (assuming it were purely a derived demand) and labour supply, while the substitution effect would suggest an increase in both. On balance, use of formal childcare would likely increase unless parents had quality or accessibility issues but labour supply effects might be less certain. For those currently paying for more than 15 (30) hours of formal childcare, the income effect would dominate, suggesting that labour supply (and childcare use, where it is a derived demand) would fall.

Parents may be constrained, however, meaning they cannot respond in the optimal way. They may not be able to find an appropriate childcare place, for example, or adjust their hours of childcare use in a flexible way. Likewise, even if they were able to optimally adjust their use of childcare, they may not be able to find a job that fits around the free hours, or adjust their hours in a flexible way (at least initially). It is for these reasons that we consider the possibility of dynamic responses to the policy, exploring whether the effects vary with time since parents initially become eligible for the subsidy.

We also need to consider that the free entitlement policies are long-standing. At the time of their child's birth, parents already know that in the absence of policy changes they will have access to a certain amount of free childcare in future. Forward-looking parents facing no constraints e.g. in terms of their ability to pay for childcare before the free entitlements kick in, or their ability to find appropriate childcare may therefore factor these entitlements into their labour supply and childcare decisions from birth. For such parents, we may thus see little or no effect of the policy at the time their child becomes eligible for free childcare (following their third birthday) or becomes eligible for more free childcare (following their fourth birthday).

If, on the other hand, parents are myopic, or face constraints which prevent them from optimising their labour supply decisions over the long-run, we might expect to see larger responses to the policy when children first become eligible, or perhaps an immediate response which then increases over time as more parents are able to respond. For the effects we observe in response to changes in free entitlement to identify the total impact of current childcare eligibility on current parental labour supply, we must assume that parents do not make labour supply decisions in anticipation of their children's future childcare eligibility. This assumption is not specific to our setting, but is common to designs exploiting birthday-based eligibility rules in contexts where those rules are known (e.g. Berlinski et al., 2011; Fitzpatrick, 2010; Goux and Maurin, 2010). In section 5.3 we provide empirical checks that validate this assumption.

4. Empirical strategies and data

Our aim is to estimate the impacts on parental labour market outcomes of children's eligibility for free part-time and full-time childcare and assess how these impacts vary with the duration of the subsidy. In this section, we describe and contrast two separate empirical strategies to recover these parameters. Both strategies exploit birthday-based eligibility rules; one uses cross-sectional data and the other uses panel data. Both strategies recover the intention-to-treat (ITT) parameters since they measure the effect of being offered free childcare rather than the effect of using free childcare, which are the relevant parameters for assessing the cost effectiveness of the policies.

4.1. Regression discontinuity (RD) approach

The first strategy we employ is a standard Regression Discontinuity (RD) design, similar to that used in Fitzpatrick (2010) and Goux and Maurin (2010). It uses point-in-time cross-sectional data and restricts attention to children born just before and just after the relevant cut-off dates in order to estimate the following models for the impact of being entitled to free full-time and part-time childcare respectively:

$$Y_i = \pi^{FT} EligFT_i + g(Days_i) + \varepsilon_i^{FT} \tag{1a}$$

$$Y_i = \pi_\tau^{PT} EligPT_{i,\tau} + f(Days_i) + \varepsilon_i^{PT} \tag{1b}$$

where Y_i is the outcome of parent of child i, $EligFT_i$ is a binary indicator for whether the child is eligible for free full-time childcare and $EligPT_{i,\tau}$ are binary indicators for whether child i is in the τth term of entitlement for free part-time childcare (where τ runs from one to four). We estimate a separate regression for each part-time treatment effect based on Eq. (1b). $g(Days_i)$ and $f(Days_i)$ are flexible functions of $Days_i$ (the running variable), the distance in days between a child's date of birth and the relevant cut-off date determining eligibility, and ε_i^{FT} and ε_i^{PT} are error terms. As the vast majority of previous studies have found parental labour supply responses concentrated largely or entirely on the entitlement to childcare of the youngest child in the household, our estimates focus on youngest children only.[19]

We estimate these regressions using 2011 UK Census data aggregated at the day of birth level.[20] The Census surveys individuals on 27 March 2011. This means that, to identify the impact of part-time childcare eligibility versus nothing, we compare the outcomes of parents of children born either side of the 1 January 2008 cut-off. In this case, children born on 31 December would be in our treatment group – towards the end of their first term of entitlement to free part-time childcare – while those born on 1 January would act as the comparator group, as they are not yet entitled to any free childcare. To provide insight into the heterogeneity of parental labour supply responses to the duration of part-time childcare entitlement we can also, in separate regressions, compare the outcomes of parents of children born either side of the 1 September 2007, 1 April 2007, and 1 January 2007 cut-offs to identify the impacts of being entitled to two, three and four terms of part-time childcare (versus one, two and three, respectively).

Our estimate of the impact of eligibility for free full-time childcare compares the outcomes of children born either side of the 1 September 2006 cut-off. This produces an estimate of the impact of eligibility for free full-time childcare 7 months (2 terms) after first gaining eligibility, as compared to being in the fourth term of entitlement to free part-time care. As in Fitzpatrick (2010), we use a parametric model to estimate the treatment parameters and control flexibly for the child's age relative to the cut-off by way of a local polynomial (quadratic) regression in $Days_i$, which we allow to have differential effects on either side of the cut-off. Our main results are based on a 90-day bandwidth and a quadratic function in age, but we present a series of robustness checks varying the size of the bandwidth and the age function.

Common to all RD designs based on Census data in the related literature, the effects we estimate are specific to children born in particular months of particular years. This is problematic if mothers of children born at different times of the year differ in unobserved ways, as suggested by the literature on seasonality of birth (Buckles and Hungerman, 2013; Clarke et al., 2019), because it would imply that the effects cannot necessarily be generalised to all children. Because researchers cannot access individual-level Census data in England that includes information on day of birth, we instead order tables of Census data aggregated at the day of birth and choose relatively large bandwidths as our preferred estimation (though we show that our estimates are robust to different bandwidths).

Similarly, as discussed in Section 2.1, there may be term-of-birth specific constraints on childcare availability affecting children which suggest the estimates from the RD approach may not represent averages across all children. Furthermore, we would like to learn how parental labour supply responses to free childcare entitlement evolve as duration of exposure increases. We cannot do this at all for eligibility for free full-time childcare, and for us to be able to interpret differences in the estimates of entitlement to free part-time childcare for children who have been eligible for different lengths of time as causal effects of the duration of entitlement on parental labour supply, we must rely on the same assumption of comparability between parents of children born close to different discontinuities as outlined above. In what follows, we propose an alternative empirical strategy that circumvents some of these challenges.

4.2. Panel data approach

The aim of the second strategy is to allow us to estimate the causal effect of free part-time and full-time childcare on parental labour supply for parents of children born in all months of the year in a way that varies with the duration of the subsidy while controlling as far as possible for unobservable differences between parents. We implement it using the Labour Force Survey (LFS), a longitudinal study following a nationally representative sample of households quarterly for up to 5 quarters. With this data set our sample size is too small to support separate RD estimates of all the termly treatment effects we are interested in. However, we have the advantage of repeated observations across individuals, which allows us to implement panel fixed effects. Moreover, we observe parents' labour supply across the whole year, allowing us to estimate the parameters of interest for parents whose children are not just born at specific times of the year.

The strategy we propose exploits two sources of variation. The first is again from birthday-based eligibility rules, as above. However, instead of focusing exclusively on parents of children born around specific cut-offs, as we do in the RD design, we compare the labour market outcomes of parents whose children are born across the whole year, conditional on age. To illustrate this, consider the effect of going from the second term of part-time childcare (PT2) to the third term of part-time childcare (PT3) as illustrated in Fig. 1. In the RD approach, we would compare the labour market outcomes of parents with children born either side of the 1 April discontinuity to estimate this effect. At a given point in time, these children would be of very similar ages. However, as Fig. 1 makes clear, conditional on, say, age = 46 months (highlighted in red in Fig. 1), it is not only children born in March and April that could be used to estimate this treatment effect. Effectively, this strategy includes all children in PT3 (i.e. children born in June, July, August, November, December, March) in our "treated" group, and all children in PT2 (i.e. children born in all other months) in our implicit "comparison" group. Similar reasoning can be applied at each age to understand which children are being compared in order to estimate each treatment effect.

These comparisons can be operationalised using the following regression:

$$Y_i = \sum_{\tau=1}^{5} \pi_\tau^{PT} EligPT_{i,\tau} + \sum_{\tau=1}^{3} \pi_\tau^{FT} EligFT_{i,\tau} + \beta' X_i + f(Age_i) + \varepsilon_i \tag{2}$$

where Y_i again is the outcome of parent of child i, $EligPT_{i,\tau}$ and $EligFT_{i,\tau}$ are binary indicators for whether child i is in the τth term of entitlement for free part-time childcare or free full-time childcare respectively. These indicators depend on the date of birth of the child and

[19] Ideally we would like to condition on the childcare eligibility of other children in the family as well but this is not possible in the Census data. We do so in our panel data approach.

[20] In the regressions, we weight each day of birth observation by the number of parents contributing to each observation.

the time of observation. τ varies from one to up to five terms for part-time care and from one to up to three terms for full-time care.[21] X_i is a vector of individual-level controls relating to the child (e.g. month of birth) and to the parent of that child (e.g. education, partnership status, ethnicity, number and age of other children). $f(Age_i)$ is a flexible function of the child's age. Standard errors are clustered at local education authority level.[22]

In Eq. (2), identification of the parameters π_τ^{PT} and π_τ^{FT} relies on $f(Age_{i,t})$ appropriately controlling for age so as not to confound the impact of entitlement to free care with the independent (generally positive) impact that children growing older has on parental labour supply. Our preferred specification for this age function includes a full set of dummies for the age in months of the youngest child in the family, and four variables measuring the number of children in the household in age bands 0–2 years, 2–4 years, 5–9 years and 10–15 years. But as we discuss in Section 6.3, our estimates are robust to alternative ways of controlling for children's ages.

Identification also relies on appropriately accounting for any differences between parents whose children are born at different points of the year. Purging the estimated values of π_τ^{PT} and π_τ^{FT} of these differences would only be possible through the inclusion of all potential confounders in the vector X_i. Instead, we make use of the fact that our dataset is longitudinal to include parent fixed effects, allowing us to account for all time-invariant differences between parents whose children are born in different months of the year. The estimating equation becomes:

$$Y_{i,t} = \sum_{\tau=1}^{5} \pi_\tau^{PT} EligPT_{i,\tau,t} + \sum_{\tau=1}^{3} \pi_\tau^{FT} EligFT_{i,\tau,t} + \beta' \widetilde{X}_{i,t} + f(Age_{i,t}) + \sigma_t + \alpha_i + \varepsilon_{i,t} \tag{3}$$

where we have added a subscript t to refer to the (calendar) time period of the observation and reflect the fact that we observe the same parent in several time periods. Here, σ_t are time effects (i.e. year or quarter dummies), α_i is an individual parent fixed effect which absorbs any time-invariant characteristics of the child and/or parent from the vector $X_{i,t}$, and any remaining time-varying variables are included in $\widetilde{X}_{i,t}$ (i.e. number and age of the children in the household, age of the mother, and year dummies).

Identification now comes from within-parent changes in labour supply over time, as their child or children move into and out of eligibility for different amounts of free childcare. For example, the mother of an only child born in August 2004 who was interviewed for the first time in November 2007 (and then every three months until November 2008) would have the treatment dummy relating to the first term of part-time entitlement (EligPT1) switched on when she was interviewed in November 2007, the treatment dummy relating to the second term of part-time entitlement (EligPT2) switched on when interviewed in February 2008, the treatment dummy for the third term of part-time entitlement switched on when interviewed in May and August 2008, and the treatment dummy relating to the first term of full-time entitlement switched on when interviewed in November 2008, with the variation arising from the fact that the child becomes older over time.

There are some similarities between our approach and the two-way fixed effects models discussed by, amongst others, Goodman-Bacon (2021), de Chaisemartin and D'Hautfœuille (2020), Callaway and Sant'Anna (2020) and Borusyak et al. (2021). Our set-up could be viewed as a two-way fixed effects model where the 'group' dimension is the childs month of birth and the 'time dimension is the child's age. However, as we show later, we find little evidence of heterogeneous treatment effects, including by term of birth (effectively a measure of 'group').[23]

When implementing the model, we make two further extensions to Eq. (3). First, in line with the literature, we allow the effect of entitlement to free childcare to differ for children who are and are not the youngest in their family by including separate eligibility indicators for youngest and non-youngest children. Second, we define our treatment variables by whether *any* child in the household is eligible for each of the different entitlements based on the time of observation and the dates of birth of all children in the household. Our estimation equation is therefore at the level of the parent and may have more than one eligibility dummy turned on, depending on the age of the children in the household. This is in order to account for the more realistic assumption that a parent's labour supply is a function of all his/her children's entitlement to childcare, rather than just an individual child. This contrasts with most of the related literature (with the exception of Lundin et al. (2008)), which instead estimates the impact of a particular (often the youngest) child's entitlement to childcare on maternal labour supply. Accordingly, as outlined above, we control for the ages of all children in the household, rather than the age of child i only.

4.3. Data and descriptive statistics

Data sources As mentioned above, our RD analysis uses the 2011 Census, which includes basic demographics, economic activity and marital status of all household members and, crucially, the birth date of all children in the household. Individual-level Census data with full date of birth are not accessible, so we order customised extracts of the data returning the number of men and women in different labour market statuses tabulated by the date of birth of their youngest child and by marital status. We obtain these data for mothers and fathers whose youngest child was born between 1 April 2003 and the 2011 Census date and use them to construct our main outcomes: the proportions of mothers/fathers whose youngest child is born on a particular day during this period who are in the labour force and the proportions who are in paid work (including self-employment).

We check that the number of births are smoothly distributed around the eligibility cut-off dates to rule out that parents might time the birth of their child to receive more free part-time or full-time child care. If so, we would see relatively more births immediately before than after the cut-off dates, invalidating the identification strategy as date of birth would be correlated with outcomes for reasons other than eligibility. Appendix Fig. A.1 plots the number of children born on each day between June 2006 and March 2008, the relevant time window for the five eligibility cut-offs we exploit. The raw numbers do not show any spikes around the cut-offs but birth patterns indicate that births are lower at weekends and during festivities and bank holidays, presumably because of a lower incidence of elective cesaereans. We also see a spiking of births on the first day of every month which could be the result of birth dates with missing day being recorded on the first day of the month by default.

[21] We are most interested in the short and medium-term impacts of extending the part-time childcare subsidy to the full-time childcare subsidy so we restrict attention to the effect of free full-time childcare in the first year of entitlement. As shown in Fig. 1, only children born between September and March contribute to estimates of the impact of four terms of part-time care and only those born between September and December to estimates of the impact of five terms of part-time care. The effects of full-time care are averages across children who were eligible for 3, 4 or 5 terms of part-time care.

[22] There are 152 Local Education Authorities (LEAs) in England. We cluster at the LEA level because LEAs are largely responsible for the local provision of education and children's social services, which could generate some correlation across the error terms of parents living in the same LEA. If the parent changes LA during the period of observation, we use the modal LEA. We have additionally estimated standard errors clustered by month of birth using a Wild cluster bootstrapping approach. The standard errors estimated in this way do not materially differ from those reported here.

[23] We also implement the checks recommended by de Chaisemartin and D'Hautfœuille (2020), finding only a very small proportion of negative weights for all eligibility dummies, and that the amount of heterogeneity required for the average treatment effect to differ from that reported here is implausibly large relative to the effect size.

We drop the first day of the month in subsequent analysis and control flexibly for day of the week and festivities by entering dummies and their interactions in the RD regressions.[24]

Our panel data approach uses the UK Labour Force Survey. Our sample includes any mother or father interviewed between 2000 and 2013 with at least one child living in the household and aged 0 to 6 at the time of the interview.[25] We drop families for whom we do not observe key characteristics, such as the date of birth of their children.[26] Table 1 provides summary statistics of key characteristics of our initial sample and our final estimation sample. The means of all the variables are very similar to each other, indicating that sampling decisions are unlikely to bias our results. Although we do not require a balanced panel, the use of parent fixed effects means that households that appear once in our sample – either because their five quarters in the LFS are left- or right- censored by our observation window, or because they attrit from the survey after their first interview – are not used. Although the exact sample size varies slightly with the outcome of interest, we end up working with a sample of about 72,000 mothers and 56,000 fathers.

We estimate Eq. (3) above for two main labour market outcomes: binary indicators for whether the mother/father is in the labour force and in paid work. We also present further results based on different measures of labour supply at the intensive margin. Specifically, we estimate the model for usual hours of work, as well as three binary indicators for working 1–15 h, 16–29 h, and 30 or more hours per week.[27] All outcomes relate to the seven days ending Sunday prior to the interview date. As LFS interviews take place continuously throughout the year, the impacts we estimate are implicitly averaged over school term-time and school holidays. Similarly, a child is defined as eligible for part-time or full-time childcare in all weeks once they reach the critical age, regardless of whether their mother is observed inside or outside school term time.

Descriptive statistics In Fig. 2, we plot the relationship between the proportion of mothers, lone mothers and fathers in the labour force (Panel A) and in employment (Panel B) and the age of the youngest child in our two samples. The labour force participation and employment rates of mothers are slightly higher in the 2011 Census than in the LFS sample (comprising years 2000–2013), reflecting the secular increase in these outcomes over time. In line with the literature, we find a positive relationship between mothers' involvement in the labour market and the age of the youngest child, with employment rates rising from 54% among mothers of 1-year-olds to 60% among 4-year-olds in the LFS and from 56% among mothers of 1-year-olds to 61% among 4-year-olds in the Census.[28] By contrast, fathers' labour force participation and employment rates do not change at all with the age of the youngest child, hovering around 95% (94%) and 91% (90%) respectively in the LFS (Census).

Employment rates of lone mothers are at least 10 percentage points below the average at all ages of the youngest child. Moreover, the relationship between labour supply and the age of the youngest child is steeper for lone mothers than for coupled mothers, which suggest that these are the sorts of mothers for whom we expect childcare affordability to be a particularly binding constraint and therefore for childcare subsidies to have a larger effect. We test whether impacts of free childcare are larger for lone mothers in both of our datasets. We now turn to our main estimates of the impact of free childcare on parental labour supply.

5. RD analysis results

Fig. 3 depicts the empirical relationship between parental labour market outcomes and the age of their youngest child. In particular, the diagrams zoom in on the five cut-offs (shown in red) at the core of our identification strategy: whether a child is born before or after January 2008, September 2007 or April 2007, January 2007, and September 2006.

The first four are, respectively, the cut-offs identifying the impact of the first term of free part-time childcare versus no free childcare, the second term of entitlement versus the first, and the third term of entitlement versus the second, and the fourth term of entitlement versus the third. The fifth cut-off is the one identifying the impact of free full-time childcare. In each figure, the dots represent the proportion of mothers/fathers in the labour force (Panels A and C, respectively) and in employment (Panels B and D, respectively) whose youngest child is of a particular age (in days) on the 2011 Census day. The superimposed curves are local polynomial regression estimates of the relationship between the outcome and the age of the youngest child (with 95% confidence bands around them), where we have estimated a different function in each of the five segments.[29]

Children to the left of the first cut-off are too young to be eligible for any free childcare at the time of the Census; children between the first and last cut-offs have been eligible for free part-time childcare for varying durations; and children to the right of the last cut-off are eligible for free full-time childcare at the time of the Census. If there were an immediate maternal labour supply response to becoming eligible for free part-time childcare, then we would expect to see a jump in outcomes at the first cut-off (between A and B). Similarly, if there were an immediate response to becoming eligible for free full-time childcare vs. free part-time childcare, then we would expect to see a jump in outcomes at the last cut-off (between E and F). We would only expect to see any discontinuous change in outcomes around the cut-offs in between if there were an effect of having been eligible for free part-time childcare for longer (e.g. for one term vs. two (C vs. B), two terms vs. three (D vs. C) or three terms vs. four (E vs. D)).

Focusing on the first cut-off, the diagram shows that, on average, having a youngest child who is eligible for the first term of free part-time childcare (versus no free childcare) has a small positive (but not statistically significant) relationship with maternal labour force participation, but no discernible effect on maternal employment, and no effect on paternal labour supply either. No effects are apparent for either mothers or fathers from the second, third and fourth terms of eligibility for free part-time care (versus the first, second and third terms respectively). Turning our attention to the last cut-off, the diagram shows that having a youngest child who is in their second term of eligibility for free

[24] We are not able to check the balance of characteristics either side of the eligibility cut-offs because we have a very limited number characteristics in the Census data available to us. Analysis by Blanden et al. (2021) based on the National Pupil Database finds that family and child characteristics around the December and March cut-off for part-time eligibility are balanced.

[25] The free part-time entitlement was fully implemented only from 2004, but we exploit the time-window from 2000 to maximise our sample size and improve the precision of the age effects. We interact all our part-time eligibility dummies with a 'pre 2004' indicator and report only the main effects of part-time eligibility estimated for the post 2004 period. In contrast, the effects of full-time eligibility are estimated on the whole period. We end our sampling period in 2013 to avoid confounding effects with the introduction of free childcare for some two-year-olds from September 2013.

[26] To be precise, in the LFS, relationships between individuals living in the same household are defined relative to the head of household. As a result, we define a respondent as a mother (father) if the head of household or spouse/cohabiting partner of the head of the household is a female (male) and if there is a child living in the household who is the head of household's natural son/daughter or step son/daughter.

[27] We choose these groupings as they relate to important thresholds used in the assessment of entitlement to in-work support in the UK and are also closely aligned with the part-time and full-time childcare offers whose effects we estimate in this paper. The outcomes relating to hours of work take a value of zero if the parent is not in work.

[28] The slightly higher participation and employment rates for 0 year olds likely reflect mothers being on maternity leave.

[29] These local polynomial estimates are calculated using an Epanechnikov kernel function.

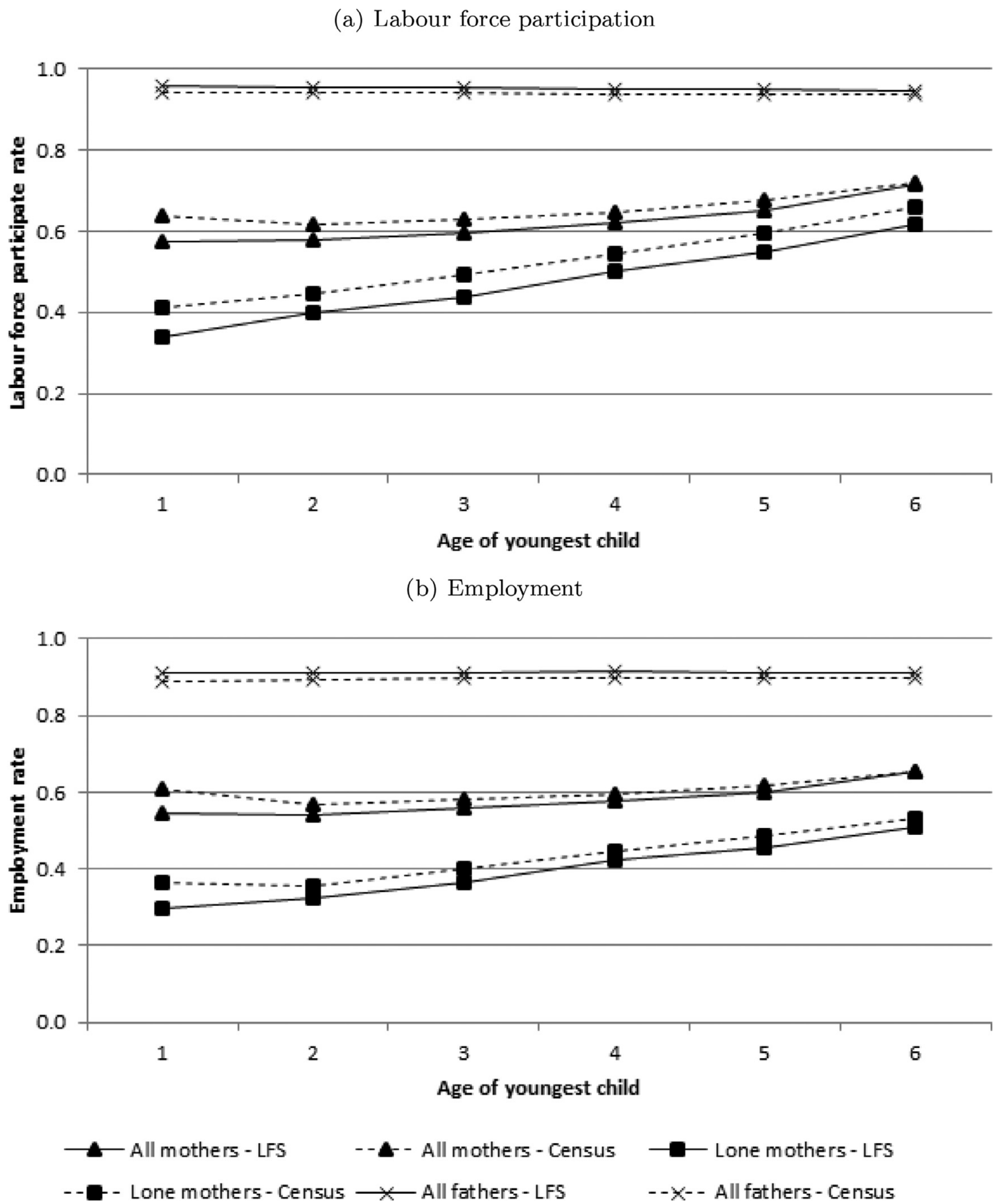

Fig. 2. Labour force participation and employment rates of parents by age of youngest child. *Notes:* These figures are based on authors' calculations based on the Labour Force Survey (LFS) 2000–2013 and the UK 2011 Census.

full-time childcare (relative to the fourth term of entitlement to free part-time childcare) is associated with a significant jump in maternal labour force participation of around 3.5 percentage points and in employment of around 1.5 percentage points. Again, for fathers, we see no jump at all.

Table 2 reports the estimates of the parametric RD regressions set out in Eqs. (1a) and (1b), where we regress the labour market outcome of interest on a treatment dummy for whether the child is to the right of the relevant cut-off, a quadratic in the child's age and an interaction between this quadratic function and the treatment dummy. The results reported in Table 2 are estimated using a 3-month bandwidth around the cut-off. In line with Fig. 2, these results show that offering free part-time childcare does not have any significant effect on the labour force participation and employment of mothers or fathers in any term after the youngest child receives a free place. However, roughly doubling the offer of free care from part-time to full-time increases the probability of mothers whose youngest child is eligible for free full-time care being in the labour force by 3.6 percentage points and the probability of being in paid work by 1.3 percentage points two terms after becoming eligible for this greater offer. We find no effect of having a youngest child entitled to free full-time childcare on fathers' labour market outcomes.

We explore the sensitivity of our results to the choice of bandwidth and to the way we control for the child's age. Appendix Table A.2 shows that estimates are relatively robust to varying the sample included in

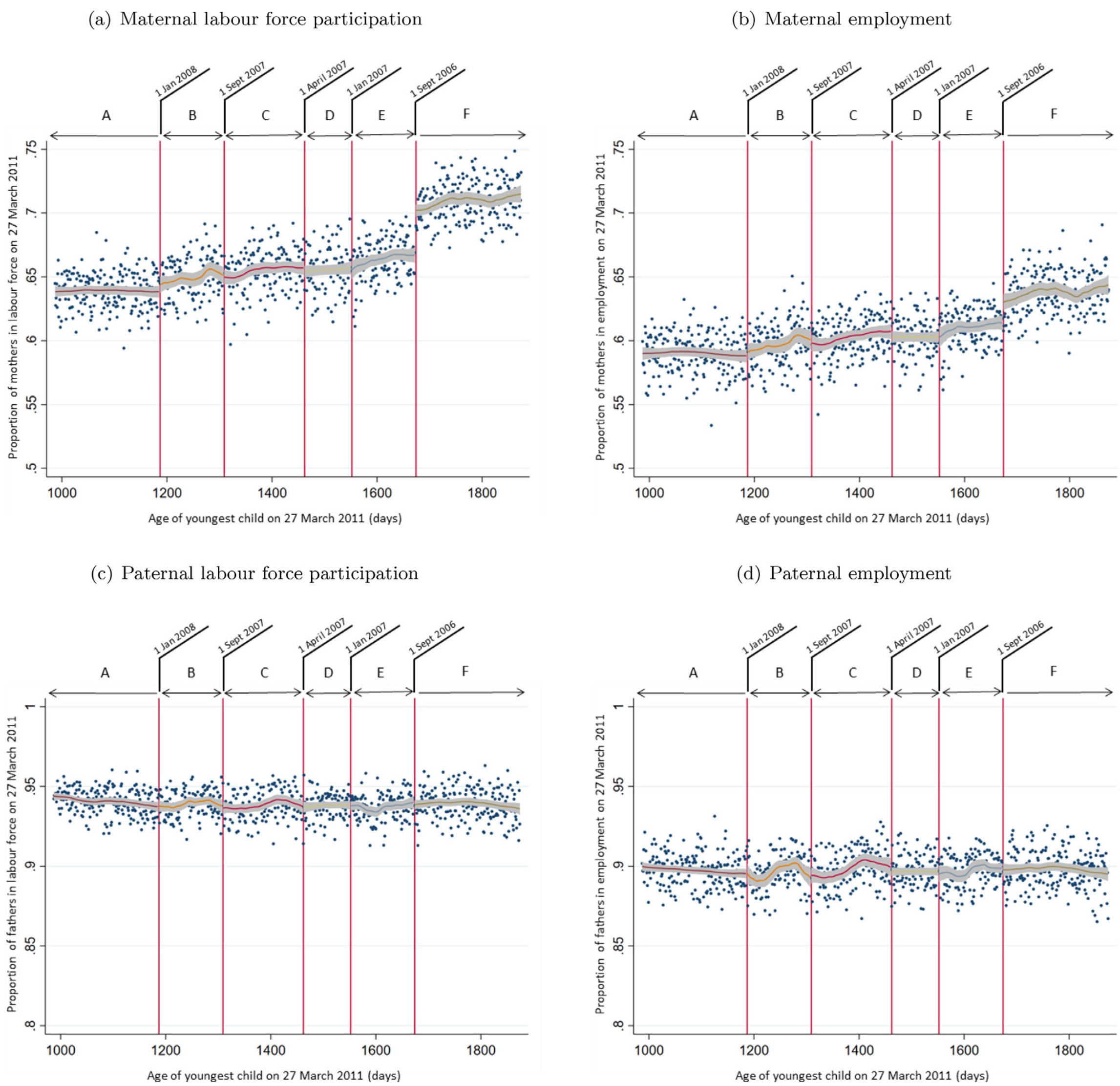

Fig. 3. Parental labour market outcomes around the cut-offs. *Notes:* Authors' calculations using the 2011 Census. Each dot depicts the outcome amongst mothers/fathers whose youngest child is of a particular age on 27 March 2011. The superimposed lines are estimates of local polynomial regressions of the outcome (on the y-axis) on the age of the youngest children on 27 March 2011. Different functions are estimated for each segment A (born after 1 Jan 2008), B (born between 1 Sept 2007 and 1 Jan 2008), C (born between 1 April 2007 and 1 Sept 2007), D (born between 1 Jan and 1 April 2007), E (born between 1 Jan 2007 and 1 Sept 2006), and F (born before 1 Sept 2006).

the estimation but somewhat sensitive to controlling for the child's age with lower and higher-order polynomials when we use smaller windows around the discontinuity.[30] In results not reported here, we have also implemented the method proposed by (Calonico et al., 2020) to choose the optimal bandwidth. Depending on the outcome and treatment effect of focus, the optimal bandwidths vary between 30 and 114 days, and the results used with those bandwidth are very similar to those presented in the paper. Further we investigate heterogeneity of impacts between lone and married mothers. While impacts appear to be slightly larger for lone mothers, differences in impacts between the two groups are statistically insignificant.

The RD results are specific to parents of children born at particular times of the year and for outcomes observed at one point in time, the Census date in 2011. We next turn to the panel data analysis to estimate effects for children born throughout the year, over a number of years, which allows us to recover termly effects for free full-time as well as free part-time care.

6. Panel data analysis

6.1. Results

Panel A of Fig. 4 graphically presents the main results of our panel data analysis of the impacts of entitlement to free part-time and full-time childcare on maternal (Panel A) and paternal (Panel B) labour force participation and employment when the youngest child in the family is eligible for these types of care. Table 3 reports the estimates underlying these figures. The first five data points on each diagram in Fig. 4 report the effect of eligibility for free part-time childcare (relative to no free childcare) and the corresponding 95% confidence intervals in each

[30] To estimate the age function, we can only use the support on either side of the birth date cut-off up to the next discontinuity. Therefore, we have to weigh up the benefit of controlling very flexibly for age, i.e. by using a higher order polynomial, with the downside of having relatively few data points to estimate a very flexible function. Appendix Table A.2 shows that in some specifications the second term of eligibility for free part-time childcare is significantly different from zero. Apart from that, the results do not differ from those shown in Table 2.

Table 2
RD estimates of the effect on parents' labour market outcomes of the youngest child's eligibility for free part-time and full-time childcare.

	(1) Mothers	(2)	(3) Fathers	(4)
	In labour force	In work	In labour force	In work
A. Part-time eligibility				
1 term PT vs nothing	0.005	0.001	0.000	−0.00
	(0.007)	(0.007)	(0.004)	(0.006)
2 terms PT vs 1 term PT	0.000	0.001	−0.000	0.003
	(0.007)	(0.008)	(0.004)	(0.005)
3 terms PT vs 2 terms PT	−0.003	−0.006	0.001	0.001
	(0.007)	(0.007)	(0.005)	(0.005)
4 terms PT vs 3 terms PT	−0.006	−0.002	0.002	−0.001
	(0.009)	(0.008)	(0.003)	(0.005)
B. Full-time eligibility relative to part-time				
2 terms FT vs 4 terms PT	0.035***	0.013*	−0.004	−0.002
	(0.006)	(0.007)	(0.004)	(0.004)

Note: This table reports estimates based on the 2011 Census from RD regressions using 3 months on each side of the relevant cut-off as bandwidth. The regression also controls for a second order polynomial in the difference between the age of the child and the relevant cut-off, an interaction between this polynomial and the cut-off, as well as day of the week dummies interacted with a dummy for whether the child was born on a holiday. All the regressions weight the observations by the number of underlying observations and use robust standard errors. We drop mothers whose youngest child was born on the first of the month from the sample; N = 177. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Fig. 4. Panel data estimates of the effect on parents' labour market outcomes. *Notes:* The coefficients plotted on these figures refer to the estimates and 95% confidence intervals of coefficients on eligibility dummies for the youngest child in equation 3 using LFS samples of mothers (fathers) with at least one child between 0 and 6 and who are observed more than once. These coefficients are estimated in a regression of a labour market outcome (indicator for labour force participation or for employment) on indicators for whether the youngest child is in a particular term of eligibility, indicators for whether any other child is in a particular term of eligibility, the number of children in the age bands 0–2; 2–4; 5–9; 10–15 in the household, age-in-month dummies of the youngest child in the household, quarter of observation dummies, whether the mother has a partner. All the regressions are linear regressions with parent-level fixed effects. The reported effect of eligibility for free part-time education is for years after 2004 (when the policy was fully in place). Standard errors are clustered at the LEA level.

Table 3
Panel data estimates of the effect on parents' labour market outcomes of the youngest child's eligibility for free part-time and full-time childcare.

	(1) Mothers	(2)	(3) Fathers	(4)
	In labour force	In work	In labour force	In work
A. Effects of part-time eligibility				
1st term	0.003	−0.004	−0.001	−0.004
	(0.004)	(0.004)	(0.003)	(0.004)
2nd term	0.011	0.000	0.001	−0.005
	(0.008)	(0.006)	(0.004)	(0.006)
3rd term	0.021**	0.008	0.003	−0.005
	(0.011)	(0.008)	(0.006)	(0.007)
4th term	0.023**	0.015	−0.001	0.000
	(0.011)	(0.009)	(0.007)	(0.009)
5th term	0.025*	0.014	0.004	−0.002
	(0.013)	(0.011)	(0.007)	(0.009)
Average effect	0.014*	0.001	0.001	−0.004
	(0.008)	(0.005)	(0.004)	(0.006)
B. Effects of full-time eligibility relative to 3rd term of part-time eligibility				
1st term	0.031***	0.011*	0.002	0.004
	(0.006)	(0.006)	(0.003)	(0.005)
2nd term	0.053***	0.026***	0.001	0.002
	(0.008)	(0.008)	(0.004)	(0.006)
3rd term	0.057***	0.035***	0.000	0.001
	(0.011)	(0.010)	(0.006)	(0.007)
Average effect	0.047***	0.024***	0.001	0.002
	(0.008)	(0.004)	(0.004)	(0.006)
C. Effects of an additional term of full-time eligibility				
2nd term FT - 1st term FT	0.022***	0.015***	−0.001	−0.002
	(0.004)	(0.004)	(0.002)	(0.003)
3rd term FT - 2nd term FT	0.004	0.009***	−0.001	−0.001
	(0.005)	(0.004)	(0.003)	(0.003)
3rd term FT - 1st term FT	0.026***	0.024***	−0.002	−0.002
	(0.007)	(0.007)	(0.004)	(0.005)
Observations	276,018		213,637	
Number of mothers/fathers	72,168		56,226	

Note: This table reports estimates and linear combinations of estimates of coefficients on eligibility dummies for the youngest child in equation 3 using LFS samples of mothers (fathers) with at least one child between 0 and 6 and who are observed more than once. These coefficients are estimated in a regression of a labour market outcome (indicator for labour force participation or for employment) on indicators for whether the youngest child is in a particular term of eligibility, indicators for whether any other child is in a particular term of eligibility, the number of children in the age bands 0–2; 2–4; 5–9; 10–15 in the household, age-in-month dummies of the youngest child in the household, quarter of observation dummies, whether the mother has a partner. All the regressions are linear regressions with parent-level fixed effects. The reported effect of eligibility for free part-time education is for years after 2004 (when the policy was fully in place). Standard errors are clustered at the LEA level. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

term of entitlement. These correspond to coefficients π_τ^{PT} for $\tau = 1..5$ in Eq. (3). These results suggest that there is little evidence that entitlement to free part-time childcare for the youngest child in the family allows more mothers to move into work. It does enable some mothers to enter the labour force, though the estimates become statistically significant only in the third term of part-time entitlement, when we estimate that eligibility for free part-time childcare increases labour force participation by 2.1 percentage points (3.4% of the baseline), with effects of a similar magnitude in the fourth and fifth terms of entitlement for parents of children who are entitled to these.

The next three data points in the diagrams show the impact of eligibility for the first, second and third terms of full-time entitlement compared to no eligibility, i.e. our estimates of the coefficients π_τ^{FT} for $\tau = 1, 2, 3$ in Eq. (3). We find significant effects on mothers whose youngest child becomes entitled to free full-time care: maternal labour force participation is 5.1 percentage points higher than without entitlement in the first term, rising to 7.8 percentage points by the 3^{rd} term. For fathers (panel B) we see no effect of the youngest child's eligibility for part- or full-time childcare on labour market participation or employment.

As discussed earlier, one innovation of this paper is our ability to assess the empirical impact of increasing entitlement to free childcare – effectively doubling the amount of free childcare available from around 3 to around 6.5 h per day – an impact whose direction is *a priori* ambiguous. Therefore, it is interesting to compare the impact of full-time eligibility to that of part-time eligibility (the relevant point estimates and standard errors are in Panel B of Table 3). Relative to the third term of part-time care, the last term in which all children are entitled to free part-time childcare, we find that increasing the childcare subsidy to cover 6.5 h a day instead of 3 increases the probability of mothers whose youngest child is eligible being in the labour force in the first term of eligibility by 3.1 percentage points. Around one third of these mothers find work, such that the probability of being in work is 1.1 percentage points higher in the first term of free full-time entitlement than in the third term of free part-time care. These effects are significant at the 1% level for labour force participation and at the 10% level for employment.

An interesting question is whether the rise in employment resulting from the entitlement to free full-time childcare is accompanied by changes in labour supply at the intensive margin. Results in Appendix Table A.3 shows that average hours worked (including the zeroes) increase by an average of 0.8 h per week by the third term of entitlement, with an increase in the proportion of mothers working 'short' part-time jobs (of 1–15 h per week) as well as full-time jobs (of at least 30 h per week). This seems to suggest that entitlement to free full-time childcare increases the hours of work of mothers with greater attachment to the labour market (who would be in work in the absence of the subsidy) and at the same time encourages some mothers to move into 'short' part-time work, though it is possible that mothers switched from zero to full-time working hours. As can be seen in Fig. 4, the impact of access to full-time childcare grows throughout the first three terms of entitlement. By the end of the first year of full-time entitlement, mothers whose youngest child is eligible are 5.7 percentage points (8.7% of the baseline) more likely to be in the labour force and 3.5 percentage points (5.9% of the baseline) more likely to be in work than in the third term of part-time eligibility. These estimates are significantly higher than those found in the first term of full-time entitlement (see Table 3, Panel B). These results suggest that it may take some time for mothers to enter the labour market and find a suitable job once their child becomes entitled to additional hours of free childcare, emphasising the importance of looking beyond the very short-term effects of childcare subsidies on labour supply.

We test for heterogeneity in these subgroups by running fully-interacted models where we interact all parameters of equation (3) with a) an indicator for mothers with a partner, b) an indicator for having low qualifications, and c) an indicator for living in an area where the unemployment rate is below median. We report these results in Appendix Table A.4.

The point estimates suggest that the effects are smaller for mothers with lower education (column a), and that the labour market participation effects are lower (but the employment effects higher) for mothers with partners (column b). None of these differences are statistically significant, however. Interestingly, we find that offering free full-time childcare has a significantly greater impact on the labour supply of mothers in lower unemployment areas (column c).

6.2. Comparison of results between the two methods

The LFS-based estimates reported in Table 3 are not exactly comparable to the Census-based estimates reported in Table 2. Indeed, the coefficients in Panel A of Table 3 refer to the effects of the first to fifth term of part-time eligibility relative to no eligibility (based on the LFS), while the coefficients in Panel A of Table 2 refer to the effects of an additional term of part-time eligibility (based on the Census). Moreover, Panel B of Table 3 refers to the effects of the first, second and third term of full-time eligibility relative to the third-term of part-time eligibility, while in the Census we can only estimate the effect of the second term of full-time eligibility relative to the fourth term of part-time eligibility. Another source of discrepancy between the results may be that the LFS results are estimated off the whole period between 2004 (2000 for full-time eligibility) and 2013, while the Census results are for 2011.

To compare the results from the two methods on more equal grounds, we estimate using LFS data the same treatment effects as are estimated in our RDD approach that is, the effect of entitlement to each additional term of part-time care, and the effect of the second term of entitlement to full-time care vs. the fourth term of entitlement to part-time care for the period 2010–13 (see Appendix Table A.5).[31] The table confirms that the results of both the RD and panel data approaches are very consistent with each other: we find little evidence of any effect of eligibility for free part-time childcare on labour force participation or employment during the first two terms of entitlement. By contrast, both approaches suggest positive effects of very similar magnitudes of entitlement to free full-time childcare (relative to free part-time childcare) on the labour supply of mothers whose youngest child is eligible (3.6 vs 3.4 ppts for labour force participation and 1.3 vs 1.5 ppts for employment).

6.3. Robustness checks

6.3.1. Anticipation effects

An important assumption underlying the interpretation of our estimates from both strategies is that parents do not change their labour supply in anticipation of their children becoming eligible for free childcare. Indeed, because the age at which free childcare is available is known to parents in advance, it is possible that their responses to the entitlement policies are affected by the future availability of care. If such responses were important, we would not be able to interpret our coefficient estimates as estimates of the policy relevant parameters. Importantly, this issue is not only relevant to our design, but to most designs exploiting birthdate-based eligibility rules in the related literature.[32]

Whether our coefficients under or over-estimate these parameters is a priori unclear, however. Parents eligible for part-time childcare may advance the take up of work in the knowledge that they will soon receive free full-time care. Alternatively, the fact that parents know they will be entitled to free full-time care later may delay their return to work because the cost of working now is higher relative to the cost of working later (for those who have no free childcare alternatives now). In the first case, our strategy would lead us to underestimate the true impact of increasing entitlement from part-time to full-time care. In the second case, it would lead us to overestimate it.

We perform two robustness checks on our panel data estimation approach to alleviate concerns about the presence of anticipation effects. To test whether mothers react to their children's future entitlement to part-time childcare we enter eligibility dummies for 2-year-olds into our model. These children are not eligible for free part-time childcare but we want to see whether mothers react in anticipation of future entitlement at this age. The second check investigates whether mothers react to future entitlement to full-time care. To do this we use data from 1992–1999, the period in time when universal free part-time care was not yet implemented, so any labour market decisions by mothers are not contaminated by universal entitlement to part-time childcare. We test whether future entitlement to full-time care has any impact on the labour force participation of mothers of 3-year-olds in this time-period.[33]

Table 4 shows the results for part-time anticipation effects in column (1) and for full-time anticipation effects in column (2). In the three terms leading up to eligibility for part-time childcare the impact on mothers' labour market behaviour are small and in opposite directions for labour force participation and employment. The point estimate on labour force participation goes up slightly throughout the three pre-entitlement terms, but none of the coefficients are not statistically significantly different from zero. Similarly, the impacts on mothers' labour supply in the (up to) 5 pre-entitlement terms in column (2), estimated using data for the years before part-time care became free for all children, show little evidence of anticipation effects. Estimates do not follow a clear pattern, are small and not statistically different from zero. This suggests that anticipation effects are unlikely to be salient during the year leading up to entitlement.[34]

[31] We do not have sufficient power to estimate treatment effects using 2011 data alone in the LFS, as a result of small sample sizes.

[32] An exception is Bauernschuster and Schlotter (2015) who use Census data for the year in which the policy was introduced in Germany. This makes it more plausible that parents respond to the childcare subsidy offer without anticipation of the policy.

[33] Some local authorities offered free part-time childcare for some children even during this time period. To account for this we control for the contemporaneous proportion of children aged 3 in free maintained nurseries at the local authority level.

[34] Of course, we cannot rule out that there may have been anticipation effects in the period of study, despite the fact that we find no evidence of anticipation effects in the 1990s.

Table 4
Robustness checks testing for the presence of anticipation effects.

	(1) Children's future entitlement to part-time childcare: effects on 2-year-olds		(2) Children's future entitlement to full-time childcare: 1992–1999 pre part-time policy years	
	In labour force	In work	In labour force	In work
A. 2-year-olds				
1st term	0.000	−0.004		
	(0.004)	(0.004)		
2nd term	0.007	−0.002		
	(0.007)	(0.006)		
3rd term	0.012	−0.001		
	(0.009)	(0.009)		
B. 3-year-olds				
1st term			0.01	0.008
			(0.006)	(0.006)
2nd term			0.009	0.001
			(0.010)	(0.010)
3rd term			0.001	−0.01
			(0.013)	(0.013)
4th term			−0.005	−0.024
			(0.016)	(0.015)
5th term			−0.015	−0.023
			(0.017)	(0.017)
Observations	267,197		200,829	

Note: The sample includes mothers with at least one child between 0 and 6 and who are observed more than once in the Quarterly Labour Force Survey. In columns (1) and (2), we report the coefficients associated with whether the youngest child in the household is in the first, second, and third term after turning 2 using observations between 2004 and 2013. In columns (3) and (4), we report the coefficients associated with whether the youngest child is in the first, second, third, fourth and fifth term after turning 3, only using observations between 1992 and 1999 (before the free entitlement policy was put in place). All the regressions are linear regressions with mother-level fixed effects. They also control for the number of children in the age bands 0–2; 2–4; 5–9; 10–15 in the household, age-in-month dummies of the youngest child in the household as well as quarter of observation dummies. Standard errors are clustered at the LEA level. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

6.3.2. Functional form of children's age effect

Controlling appropriately for the age of the youngest child and the age of any other children in the household is crucial to isolate the effect of the policy on maternal labour supply in an unbiased fashion. Our main specification controls for the age of the youngest child through age-in-month dummies and for the number and age of other children in the household through a set of variables measuring the number of children in the following age bands: 0–2; 2–4; 5–9; 10–15. We investigated the sensitivity of our results to controlling for the ages of all children in the household in a variety of alternative ways. Table 5 reports the results of three such specifications. In Column (1) we add cubic controls for the age in days of up to the next six youngest children in the household (in addition to our baseline age controls). Column (2) displays results when adding age in month dummies for the second youngest child to our baseline age controls, and Column (3) when adding age in month dummies also for the third youngest child. Looking across these models, estimates of the impact of entitlement to free part-time and full-time childcare remain remarkably stable and are almost identical to the main results reported in Table 4, reassuring us that age effects are not driving our results.

6.3.3. Capacity constraints

Our next robustness check tests whether our results are affected by capacity constraints, which, as discussed above, may be a particular problem for our estimates of the effect of entitlement to free part-time childcare. In Section 2, we discussed that these constraints might arise in two ways and we now present the results of two robustness checks we perform to investigate whether these capacity constraints weaken the labour supply responses of parents to the free part-time childcare offer.

The first reason capacity constraints may arise is if children born in different terms of the year face differential chances of securing a place at nursery. Because nursery places in England tend to become available from September, this could weaken the labour supply responses of parents of autumn- and spring-born children relative to those of summer-born children. We investigate whether this is the case by estimating a very flexible specification in which we allow the impact of each term of eligibility for part-time care to vary with the child's term of birth and then test whether the impacts are equal across all terms of birth. We report these estimates in the first three columns of Table 6 and the p-value of the tests in the fourth column. Results show that we cannot reject that the impact of each term of entitlement is the same across mothers whose youngest child is born in different terms, suggesting that this type of capacity constraint is not leading us to underestimate the effect of entitlement to part-time childcare on maternal labour supply.

The second reason capacity constraints could affect the impact of free part-time childcare is that, although places should have been universally available from 2004, full coverage of funded places was not achieved until about 2007, though this varied a lot across areas. To check whether our estimates of the impact of entitlement to free part-time childcare might be downward-biased, we add controls for the availability of funded places in the mothers' local authority of residence (as the proportion of 3-year-olds with a funded place in the local authority). These results are reported in the fifth column of Table 6. The estimated impacts of entitlement to part-time care are very similar for labour force participation and the probability of being in work to those in our baseline specification, again suggesting that capacity constraints are not significantly downward-biasing our estimates of the impact of entitlement to part-time care.[35]

[35] We have also run a specification interacting the proportion of 3-year-olds with a funded place with the treatment effects for entitlement to each term of part-time care. These results suggest that while the effects may be slightly larger for areas with full or close to full capacity of funded places, because coverage is

Table 5
Robustness checks testing the sensitivity of the results to different ways of controlling for children's age.

	(1) With cubics in age for all children		(2) With age in months dummies for 2nd youngest child		(3) With age in months dummies for 2nd and 3rd youngest child	
	Labour force	In work	Labour force	In work	Labour force	In work
A. Effects of part-time eligibility						
1st term	0.003	−0.004	0.003	−0.004	0.003	−0.004
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
2nd term	0.011	0.000	0.011	0.000	0.010	−0.001
	(0.008)	(0.006)	(0.008)	(0.006)	(0.008)	(0.006)
3rd term	0.021**	0.008	0.022**	0.008	0.022**	0.008
	(0.011)	(0.008)	(0.010)	(0.008)	(0.011)	(0.008)
4th term	0.023**	0.015	0.024**	0.015	0.023**	0.015
	(0.011)	(0.009)	(0.011)	(0.009)	(0.012)	(0.010)
5th term	0.025*	0.014	0.025**	0.014	0.025*	0.014
	(0.013)	(0.011)	(0.013)	(0.011)	(0.013)	(0.011)
B. Effects of full-time eligibility relative to 3rd term of part-time eligibility						
1st term	0.031***	0.011*	0.031***	0.011*	0.031***	0.011*
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
2nd term	0.053***	0.026***	0.053***	0.026***	0.053***	0.026***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
3rd term	0.057***	0.035***	0.057***	0.035***	0.057***	0.035***
	(0.011)	(0.010)	(0.011)	(0.010)	(0.011)	(0.010)
Observations			276,018			

Note: This table reports estimates of the same models as those reported in Panels A and B of Table 3 (for mothers) but where we control for children's age in different ways. Results in column (1) control for children's age by using age bands and including a cubic in the age in days of up to six youngest children in the family. Results in column (2) use age bands and control for the age of the two youngest children using age in months dummies. Results in column (3) add age in months dummies for the second and third youngest child. $^{*}p < 0.10$, $^{**}p < 0.05$, ***p< 0.01.

Table 6
Robustness checks testing for the effect of capacity constraints on the estimated effects of the youngest child's entitlement to free part-time childcare.

	(1)	(2)	(3)	(4)	(5)
	Allowing for term of birth specific coefficients				Controlling for funded places
	Spring borns	Summer borns	Autumn borns	*p-value*	(all terms of birth pooled)
A. Dependent variable: Mother is in the labour force					
1st term PT eligibility	0.007	0.004	0.000	0.585	0.003
	(0.006)	(0.006)	(0.006)		(0.005)
2nd term PT eligibility	0.016*	0.011	0.009	0.630	0.009
	(0.009)	(0.009)	(0.009)		(0.008)
3rd term PT eligibility	0.030***	0.020*	0.021*	0.360	0.020*
	(0.011)	(0.011)	(0.011)		(0.011)
4th term PT eligibility	0.027**		0.022*	0.531	0.023**
	(0.012)		(0.012)		(0.011)
5th term PT eligibility			0.024*		0.025**
			(0.013)		(0.013)
P-value (joint equality across all terms)			0.879		
B. Dependent variable: Mother is employed					
1st term PT eligibility	0.001	−0.008	−0.001	0.139	−0.004
	(0.006)	(0.005)	(0.005)		(0.004)
2nd term PT eligibility	0.004	−0.006	0.002	0.260	−0.001
	(0.007)	(0.008)	(0.007)		(0.006)
3rd term PT eligibility	0.012	0.004	0.008	0.568	0.007
	(0.010)	(0.009)	(0.009)		(0.008)
4th term PT eligibility	0.020**		0.011	0.125	0.015
	(0.010)		(0.010)		(0.010)
5th term PT eligibility			0.012	0.014	
			(0.011)		(0.011)
P-value (joint equality across all terms)				0.407	
Observations		276,018			271,339

Note: This table reports estimates of the same models as those reported in Panels A of Table 3 (for mothers) with the following differences: in the specification reported in columns (1) to (3), we also interact the eligibility dummies pertaining to the youngest child with dummies for his/her term of birth; in the specification reported in column (5), we also control for the proportion of 3 year olds in the LEA of residence that have a funded part-time nursery place. In column (4), we report the p-value of a test of equality across the coefficients reported in the first three columns. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

7. Discussion of the results and analysis of childcare use

The results presented so far suggest that there is little impact of entitlement to free part-time care on the labour supply of either mothers or fathers, but larger impacts of moving from part-time to full-time care for mothers whose youngest child becomes eligible. In relation to the literature, our estimates of the impact of free part-time childcare are lower than the positive and significant impacts of similar policies found by Bauernschuster and Schlotter (2015) in Germany and Berlinski and Galiani (2007) in Argentina. In comparison with countries where free or highly subsidised childcare is offered full-time, our estimates imply that the impact of free full-time childcare in England is roughly similar to those found in Spain (Nollenberger and Rodriguez-Planas, 2015), thus standing in between the very small impacts found in Norway in the late 1970s (Havnes, Mogstad, 2011) and in the US in the early 2000s (Fitzpatrick, 2010) and the large impacts found in Quebec (Baker et al., 2008). So while our estimates suggest that free full-time childcare is more effective at increasing maternal labour supply than free part-time childcare, it cannot be said to have dramatically transformed mothers' labour market outcomes over this period.

There are at least four reasons why the free part-time and full-time childcare policies that we have studied may not have been more effective at increasing parental labour supply. First, the maternal employment rate was hovering around 57% when free part-time childcare was introduced in England in the early 2000s. In contrast, when free part-time childcare was introduced in Argentina and Germany, the employment rate of mothers with 3 and 4 year olds was lower than in England (around 40% in Argentina and 50% in Germany). The labour supply decisions of mothers at the margin may thus have been more difficult to affect in our context.

The second reason why the childcare policies we study may not have had larger impacts on parental labour supply is that the offer of free childcare may not start early enough following their child's birth to prevent mothers from leaving their jobs and detaching from the labour force. In contrast to Quebec, where subsidised full-time childcare is offered to children aged 0 to 5, in England the universal entitlement to a free part-time childcare place starts at age 3 and children do not start school (and hence become entitled to a free full-time childcare place) until age 4. While low- and middle- income working families benefit from other forms of childcare support during this critical early period, these subsidies may not be high enough to incentivise mothers, especially low-income mothers, to return to work quickly after their child's birth (Blundell et al., 2016).

Third, the childcare on offer in England may not be sufficiently generous or sufficiently flexible to enable parents to work. In Quebec, for example, parents could access up to 10 h of subsidised childcare per day, while the offer of free full-time childcare that we have analysed is for 6.5 h a day that can only be taken at set times. Certainly, the fact that there is no free entitlement to childcare for parents outside school term time places a significant constraint on the policy's ability to remove financial barriers to work, which may be particularly disadvantageous for lone parents or those from less educated backgrounds.

Finally, since the late 1990s, England has experienced a large expansion of its private childcare market, and the rate of both formal and informal childcare use was already high, especially amongst working families, where over 80% of 3 and 4 year olds used formal childcare and over 40% of 3 and 4 year olds used informal childcare (Bryson et al., 2012). In this context, it is possible that there was limited scope to increase the use of childcare overall, thus not freeing much time for parents desiring to work to enter the labour force.

We investigate this issue further by using another dataset, the Family Resources Survey (FRS), which contains detailed information about households' use of different types of childcare (both formal and informal). Specifically, we use repeated cross-sections from the FRS to estimate the effect of eligibility for free part-time and full-time childcare on measures of childcare use at the child level.[36] The cross-sectional nature of the FRS necessitates that we estimate a version of equation (3) in which we do not include child (or mother) fixed effects but instead include a rich vector of time-invariant characteristics that would be dropped in a fixed effects specification:

$$C_{i,t} = \sum_{\tau=1}^{5} \gamma_{\tau}^{PT} EligPT_{i,t,\tau} + \sum_{\tau=1}^{3} \gamma_{\tau}^{FT} EligFT_{i,t,\tau} + \beta' X_{i,t} + f(Age_{i,t}) + \sigma_t + \varepsilon_{i,t} \quad (4)$$

where $C_{i,t}$ is a variable measuring use of childcare by child i observed at time t, the vector $X_{i,t}$ includes a set of permanent and time-varying characteristics about the parent(s) and children in the household,[37] age is controlled for in age in month dummies of the child and other covariates are the same as those used for our main analysis. In this specification, the impacts of eligibility rules on childcare use will be causal under the fairly strong assumption that there are no unobserved systematic differences between parents of children born in different terms of the year. We therefore refrain from giving a strong causal interpretation to these results.[38]

We estimate Eq. (4) for different types of childcare use, each measured as whether a child accesses that type of childcare and the number of hours per week of each type of care used. We focus on any type of care provided outside the immediate family such as parents or primary caregivers (any care). We distinguish between subsidisable care (i.e. care provided by the sorts of establishments where parents can take up their entitlement to free part-time childcare[39]), and informal care (i.e. time spent being cared for by family members other than immediate family, e.g. by grandparents, friends, unregistered childminders or nannies). Appendix Table A.1 summarises how these outcome variables vary by the age of the youngest child.

Table 7 reports our estimates of Eq. (4) for the youngest child in the family.[40] The top panel displays the impact of eligibility for free part-time childcare in the first to fifth terms of entitlement relative to no eligibility. The bottom panel displays the impact of eligibility for free full-time childcare in the first to third terms of entitlement relative to the third term of part-time entitlement. Column (3) provides strong evidence that becoming eligible for free part-time childcare increases the likelihood of using subsidisable care, and that this likelihood rises further when a child becomes eligible for free full-time childcare. Specifically, the use of subsidisable care increases by 12 percentage points by

high throughout our period of observation, the overall effects remain very close to the main effects reported in Table 3.

[36] The Family Resources Survey (DWP, 2016) is a yearly repeated cross-sectional household survey that collects information on the incomes and circumstances of private households in the UK. Our sample includes children between the age of 2 and 5.5 at the time of the interview, living in families in England who are interviewed between April 2005 and March 2013. The FRS collects detailed information on all the ways in which children are looked after in a reference week.

[37] These include the age and educational qualifications of the main carer and (if present) her partner, an indicator for whether the mother is married or cohabiting, a dummy for whether the child has any siblings, local authority dummies, a dummy indicating whether the local authority of residence operated a school admission policy in which all children start full-time education in the September after they turn four and month of birth indicators of the youngest child.

[38] We explored the possibility of using pseudo-cohort methods to allow for mother- or child-level fixed effects in equation (4), which would have allowed us to present a two-sample two stage least squares estimate of the causal impact of childcare use on maternal labour supply. Unfortunately, the sample size of the FRS is too small to implement such a method.

[39] These will typically be day nurseries and also state-run infant or primary schools.

[40] Results focusing on all children are available upon request and are similar to those for youngest children.

Table 7
Estimates of the effects of the youngest child's eligibility for free part-time and full-time childcare on the youngest child's childcare use.

	(1) Any care	(2)	(3) Subsidisable care	(4)	(5) Informal care	(6)
	Any use	Weekly hours	Any use	Weekly hours	Any use	Weekly hours
A. Effects of part-time eligibility						
1st term	−0.0473	−2.203	0.0809**	1.643*	−0.155***	−3.419***
	(0.0319)	(1.447)	(0.0334)	(0.851)	(0.0411)	(1.064)
2nd term	−0.0713*	−3.375*	0.0651	0.527	−0.176***	−3.314**
	(0.0427)	(2.040)	(0.0455)	(1.291)	(0.0553)	(1.407)
3rd term	−0.0761	−2.679	0.117**	1.146	−0.183**	−3.784**
	(0.0498)	(2.544)	(0.051)	(1.608)	(0.0752)	(1.897)
4th term	−0.0579	0.289	0.177***	2.141	−0.127	−0.591
	(0.0655)	(3.058)	(0.0631)	(2.169)	(0.0881)	(2.104)
5th term	−0.0697	−0.191	0.163**	1.940	−0.139	−0.895
	(0.0688)	(3.643)	(0.0726)	(2.539)	(0.0994)	(2.490)
Average effect	−0.0644	−2.06	0.108**	1.345	−0.162***	−2.847*
	(0.0428)	(2.059)	(0.0439)	(1.328)	(0.0592)	(1.478)
B. Effects of full-time eligibility relative to 3rd term of part-time eligibility						
1st term	0.0142	1.908	0.0814**	1.867	0.0034	1.356
	(0.0329)	(1.948)	(0.0356)	(1.499)	(0.0474)	(1.049)
2nd term	0.0428	3.648	0.0923**	2.770	0.0537	2.092
	(0.0394)	(2.286)	(0.0416)	(1.690)	(0.0585)	(1.304)
3rd term	0.0244	4.796*	0.11**	3.713**	0.0137	2.21
	(0.0401)	(2.531)	(0.0446)	(1.759)	(0.0640)	(1.431)
Average effect	0.0257	3.513	0.0956**	2.839*	0.0205	1.888
	(0.0359)	(2.166)	(0.0393)	(1.593)	(0.0532)	(1.171)
Observations				11,187		

Note: The sample is children aged 2 to 5.5 at the time of the interview, living in families in England interviewed between April 2005 and March 2013 in the Family Resources Survey (FRS). We include different eligibility dummies for the youngest child and other children, and only report here the ones for the youngest child. All the regressions are linear regressions and they also control for the age of the child in month dummies, child's month of birth dummies, age and educational qualifications of the main carer and partner (if present), an indicator for whether the mother is married or cohabiting, a dummy for whether the child is the only child, Local Authority dummies, and a dummy indicating whether the Local Authority of residence operated a school admission policy whereby children start school the September after they turn 4. We also control for the age of other children in the household in the age bands 0–2; 2–4; 5–9; 10–15 in the household. Standard errors are clustered at the LEA level. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

the third term of part-time eligibility and increases by another 11 percentage points by the third term of full-time eligibility. In other words, the policies that we study have some "bite" in increasing the use of the types of childcare they subsidise. However, there is little evidence that this rise in the use of subsidisable care means that children are spending more time in childcare overall: Columns (1) and (2) show that there is no change in the likelihood of using any form of childcare outside the immediate family in response to the offer of free part-time or full-time childcare, and only a small increase in the number of hours used per week when children become entitled to free full-time childcare. This suggests that there is substantial crowding-out of other forms of care by free formal childcare arrangements. As Columns (5) and (6) indicate, parents primarily substitute away from informal care arrangements when formal care becomes free of charge, especially during the first three terms of part-time entitlement.[41]

8. Conclusion

As many countries are considering increasing the number of hours of free or highly subsidised childcare available to families with pre-school children, it is important to understand the impacts that such extensions are likely to have on parental labour supply. In the past decade, many studies have estimated the impact of free or subsidised part-time or full-time childcare on maternal labour supply in various contexts and using different methods. Our paper contributes to this literature by estimating the impact of extending the offer of free childcare from half day to the whole of the school day. In doing so, it also provides the first evaluation of these two major policies on the labour supply of all parents in England.

Our estimates from both the RD and panel data approaches suggest that there is little impact of entitlement to free part-time childcare on the labour supply of either mothers or fathers, but larger and significant impacts of moving from part-time to full-time care for mothers whose youngest child becomes eligible. Panel data estimates show that in the first term of full-time entitlement, the probability of being in the labour force is 3.1 percentage points higher and the probability of being in work is 1.2 percentage points higher than in the third term of free part-time entitlement. These impacts increase in the months following initial entitlement, so that by the end of the first year, mothers whose youngest child is eligible for free full-time care are 5.7 percentage points more likely to be in the labour force and 3.5 percentage points more likely to be in work than mothers whose youngest child is eligible for free part-time care. Our estimates based on Census data are in line with these results.

When free part-time childcare was introduced in England in the early 2000s, the maternal employment rate was hovering around 57%. England was experiencing a large expansion of its private childcare market, and the rate of formal and informal care was high, especially amongst working families. In this context, it is perhaps unsurprising that providing 2.5 or 3 h a day of free childcare was too weak an incentive to encourage many new mothers to join the labour force.

Viewed across the entire observation period, the part-time entitlement did allow a few mothers already in work to switch from part-time to full-time work, but for most, the policy acted as an income transfer that families used to substitute away from informal care and/or reduce their out-of-pocket expenses on formal care without substantially affecting their labour supply. There are also other factors that may have contributed to the relatively small impacts that we find. The offer of

[41] In related work Yamaguchi et al. (2018) find that formal childcare crowds out informal care particularly among mothers with a strong labour market attachment, explaining low impacts of childcare expansion in Japan.

free childcare may not start early enough following their child's birth to prevent mothers from leaving their jobs and detaching from the labour force. It may also be the case that the offer may not be sufficiently generous or sufficiently flexible to enable parents to work. Finally, because any part-time arrangement is due to be temporary (until the start of school), mothers may not be willing to rethink their participation decisions while the current childcare arrangement is only in place for a few terms.

In considering whether to extend childcare subsidies, there are obviously trade-offs in terms of how the government should spend its limited resources. Offering more hours per week or more weeks per year for all children would either increase the total cost of the policy or necessitate a reduction in funding per child, potentially compromising the quality of provision that could be accessed, with consequences for child development. Governments may therefore wish to consider offering more support to a smaller number of parents – rather than less support to all parents – in order to maximise the cost effectiveness of childcare subsidies.

Appendix

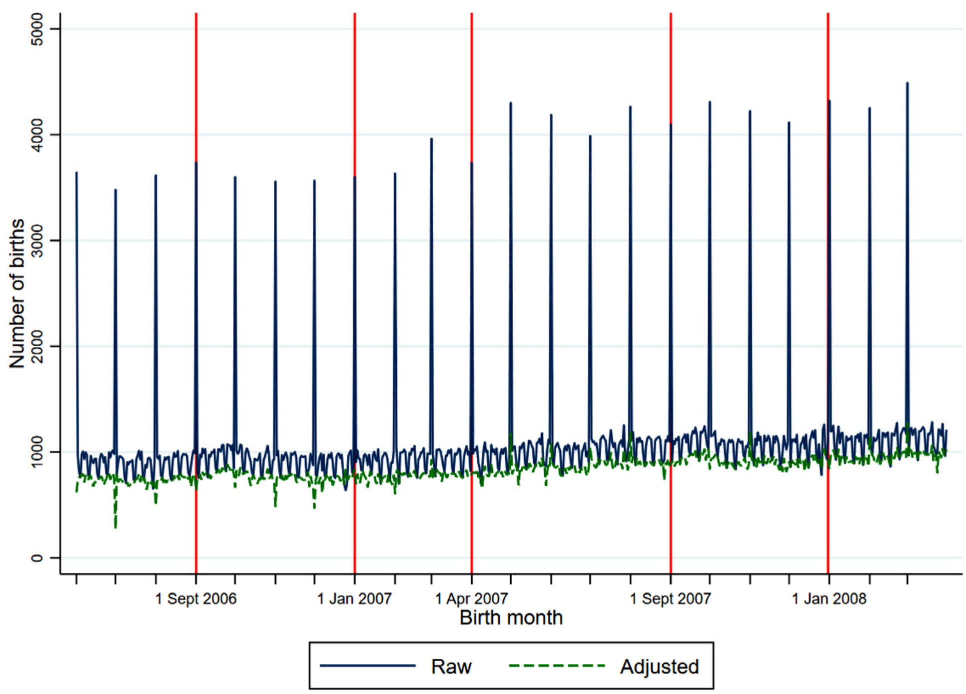

Fig. A.1. Distribution of births around the cut-offs. *Notes:* Authors' calculations using the 2011 Census. The blue line represents the total (raw) number of children born on each day before or after the relevant cut-offs, labelled by a red vertical line. The green line represents the residual of a regression of number of births on separate dummies for days of the week dummies, bank holidays/festivities, the first day of the month and interactions between of all of these variables. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

Table A.1
Average childcare use by age of youngest child.

Age of youngest child:	0	1	2	3	4	5
Any childcare						
Any use	0.380	0.620	0.710	0.820	0.900	0.890
	(0.007)	(0.008)	(0.008)	(0.007)	(0.006)	(0.007)
Weekly hours	5.381	12.711	14.785	19.096	30.292	31.428
	(0.174)	(0.254)	(0.285)	(0.317)	(0.413)	(0.413)
Formal, subsidisable care						
Any use	0.106	0.302	0.476	0.696	0.808	0.788
	(0.005)	(0.007)	(0.008)	(0.009)	(0.008)	(0.009)
Weekly hours	1.142	4.656	6.642	11.273	22.739	23.587
	(0.079)	(0.158)	(0.185)	(0.216)	(0.340)	(0.288)
Formal, non-subsidisable care						
Any use	0.03	0.076	0.087	0.085	0.133	0.193
	(0.003)	(0.004)	(0.005)	(0.005)	(0.007)	(0.008)
Weekly hours	0.553	1.638	1.932	1.335	1.597	1.363
	(0.060)	(0.105)	(0.127)	(0.105)	(0.114)	(0.092)
Informal care						
Any use	0.300	0.430	0.430	0.420	0.410	0.410
	(0.007)	(0.008)	(0.008)	(0.007)	(0.006)	(0.007)
Weekly hours	3.686	6.417	6.210	6.488	5.955	6.478
	(0.141)	(0.187)	(0.199)	(0.226)	(0.236)	(0.302)

Note: This table reports the means and standard errors of the means (in parenthesis) of different measures of childcare use for the youngest child in a sample of families with at least a child between 0 and 5.5 years old living in England and observed in the Family Resources Survey (FRS) for 2005–2013. The sample size is 19,565. Subsidisable care includes day nurseries, infant and primary schools; informal care includes unregistered childminders, friends and non-parental relatives. The final category of "formal, non-subsidisable care" is not shown here.

Table A.2
Sensitivity of the RD estimates to the choice of the bandwidth and age function specification.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Bandwidth size:	1 months around the cut-off			2 months around the cut-off			3 months around the cut-off		
	Polynomial degree of age fct.			Polynomial degree of age fct.			Polynomial degree of age fct.		
	1	2	3	1	2	3	1	2	3
	A. Dependent variable: Maternal labour force participation								
1 term PT vs nothing	−0.003	−0.013	0.009	0.001	−0.000	−0.010	0.005	0.005	−0.010
	(0.008)	(0.012)	(0.017)	(0.006)	(0.008)	(0.011)	(0.005)	(0.007)	(0.009)
2 terms PT vs 1 term PT	0.007	0.034***	0.037**	−0.009	0.019**	0.027***	−0.005	0.000	0.015*
	(0.007)	(0.010)	(0.014)	(0.006)	(0.008)	(0.010)	(0.005)	(0.007)	(0.009)
3 terms PT vs 2 terms PT	−0.002	0.005	−0.029	−0.004	0.001	−0.003	−0.005	−0.003	0.003
	(0.008)	(0.016)	(0.019)	(0.006)	(0.009)	(0.013)	(0.005)	(0.007)	(0.010)
4 terms PT vs 3 terms PT	−0.007	−0.013	−0.031	−0.001	−0.004	−0.018	−0.000	−0.006	−0.002
	(0.011)	(0.018)	(0.022)	(0.007)	(0.012)	(0.016)	(0.005)	(0.009)	(0.013)
2 terms FT vs 4 terms PT	0.041***	0.034***	0.023	0.031***	0.039***	0.039***	0.031***	0.035***	0.039***
	(0.007)	(0.010)	(0.014)	(0.005)	(0.007)	(0.008)	(0.004)	(0.006)	(0.007)
	B. Dependent variable: Maternal employment								
1 term PT vs nothing	−0.009	−0.016	−0.000	−0.001	−0.004	−0.015	0.001	0.001	−0.012
	(0.008)	(0.011)	(0.016)	(0.006)	(0.009)	(0.012)	(0.005)	(0.007)	(0.009)
2 terms PT vs 1 term PT	0.004	0.032***	0.034**	−0.008	0.018**	0.024**	−0.007	0.001	0.013
	(0.007)	(0.010)	(0.015)	(0.006)	(0.009)	(0.011)	(0.005)	(0.008)	(0.010)
3 terms PT vs 2 terms PT	−0.002	−0.004	−0.035**	−0.007	−0.002	−0.006	−0.006	−0.006	−0.001
	(0.008)	(0.014)	(0.017)	(0.006)	(0.009)	(0.011)	(0.005)	(0.007)	(0.010)
4 terms PT vs 3 terms PT	−0.003	−0.004	−0.014	0.003	−0.002	−0.010	0.005	−0.002	0.001
	(0.012)	(0.017)	(0.021)	(0.007)	(0.012)	(0.016)	(0.005)	(0.008)	(0.012)
2 terms FT vs 4 terms PT	0.016**	0.007	0.000	0.012**	0.014*	0.011	0.014***	0.013*	0.012
	(0.008)	(0.011)	(0.016)	(0.005)	(0.008)	(0.010)	(0.005)	(0.007)	(0.009)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Bandwidth size:	1 months around the cut-off			2 months around the cut-off			3 months around the cut-off		
	C. Dependent variable: Paternal labour force participation								
1 term PT vs nothing	0.001	0.002	−0.002	−0.001	0.003	0.001	−0.001	0.000	0.003
	(0.005)	(0.008)	(0.010)	(0.003)	(0.005)	(0.006)	(0.003)	(0.004)	(0.005)
2 terms PT vs 1 term PT	0.002	−0.003	0.001	−0.002	0.000	0.005	−0.005*	−0.000	0.001
	(0.005)	(0.009)	(0.014)	(0.003)	(0.006)	(0.008)	(0.003)	(0.004)	(0.006)
3 terms PT vs 2 terms PT	−0.000	0.002	−0.007	−0.001	0.000	0.001	−0.003	0.001	0.000
	(0.005)	(0.009)	(0.013)	(0.004)	(0.006)	(0.009)	(0.003)	(0.005)	(0.006)
4 terms PT vs 3 terms PT	−0.002	0.001	−0.005	−0.001	0.004	−0.003	−0.003	0.002	0.002
	(0.004)	(0.006)	(0.008)	(0.003)	(0.004)	(0.006)	(0.002)	(0.003)	(0.005)
2 terms FT vs 4 terms PT	−0.010**	−0.014**	−0.004	−0.004	−0.007*	−0.017***	−0.004	−0.004	−0.010**
	(0.004)	(0.006)	(0.008)	(0.003)	(0.004)	(0.006)	(0.003)	(0.004)	(0.004)
	D. Dependent variable: Paternal employment								
1 term PT vs nothing	0.002	0.009	0.003	−0.002	0.004	0.005	−0.004	−0.001	0.008
	(0.007)	(0.011)	(0.015)	(0.004)	(0.007)	(0.010)	(0.004)	(0.006)	(0.008)
2 terms PT vs 1 term PT	0.005	−0.001	0.001	−0.002	0.004	0.007	−0.009***	0.003	0.004
	(0.006)	(0.010)	(0.014)	(0.004)	(0.006)	(0.009)	(0.003)	(0.005)	(0.007)
3 terms PT vs 2 terms PT	−0.002	0.002	−0.001	−0.001	−0.000	−0.001	−0.005	0.001	−0.001
	(0.006)	(0.010)	(0.015)	(0.005)	(0.006)	(0.010)	(0.004)	(0.005)	(0.007)
4 terms PT vs 3 terms PT	−0.007	−0.004	−0.005	−0.003	−0.001	−0.008	−0.004	−0.001	−0.003
	(0.006)	(0.009)	(0.012)	(0.004)	(0.006)	(0.009)	(0.003)	(0.005)	(0.007)
2 terms FT vs 4 terms PT	−0.008*	−0.015**	−0.008	−0.003	−0.003	−0.016**	−0.005*	−0.002	−0.005
	(0.004)	(0.006)	(0.008)	(0.004)	(0.005)	(0.006)	(0.003)	(0.004)	(0.005)

Note: This table reports estimates based on the 2011 Census from RD regressions varying bandwidth size and the degree of the polynomial function used to control for the youngest child's age (in days). All the regressions weight the observations by the number of underlying observations and use robust standard errors. The regression also controls for day of the week dummies interacted with a dummy for whether the child was born on a holiday. We drop mothers whose youngest child was born on the first of the month from the sample. In columns (1)-(3), N = 62; in columns (4)-(6), N = 120; in columns (7)-(9), N = 177. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.3
Panel data estimates of the effect on mothers' working hours of the youngest child's eligibility to free part-time and full-time childcare.

	(1) Weekly working hours	(2) ≥0 & <16 h	(3) ≥16 & <30 h	(4) 30+ h
A. Effects of part-time eligibility				
1st term	−0.057	−0.004	−0.004	0.004
	(0.123)	(0.004)	(0.006)	(0.004)
2nd term	−0.024	0.003	−0.01	0.007
	(0.193)	(0.006)	(0.008)	(0.006)
3rd term	0.035	0.010	−0.010	0.009
	(0.273)	(0.008)	(0.010)	(0.008)
4th term	0.297	0.016*	−0.017	0.019**
	(0.318)	(0.009)	(0.013)	(0.009)
5th term	0.344	0.013	−0.018	0.022**
	(0.373)	(0.011)	(0.014)	(0.011)
Average effect	0.061	0.006	−0.01	0.01***
	(0.105)	(0.006)	(0.009)	(0.003)
B. Effects of full-time eligibility relative to 3rd term of part-time eligibility				
1st term	0.318	0.008	−0.005	0.009
	(0.194)	(0.006)	(0.007)	(0.006)
2nd term	0.600**	0.011	0.003	0.013*
	(0.273)	(0.008)	(0.009)	(0.007)
3rd term	0.838***	0.012	0.003	0.020**
	(0.337)	(0.009)	(0.012)	(0.009)
Average effect	0.600**	0.0100*	0.000	0.015***
	(0.264)	(0.006)	(0.005)	(0.003)
C. Effects of an additional term of full-time eligibility				
2nd term FT - 1st term FT	0.282***	0.002	0.008	0.004
	(0.115)	(0.004)	(0.005)	(0.004)
3rd term FT - 2nd term FT	0.238**	0.002	0.000	0.006**
	(0.117)	(0.004)	(0.005)	(0.003)
3rd term FT - 1st term FT	0.520***	0.004	0.008	0.010
	(0.198)	(0.006)	(0.008)	(0.006)
Observations	273,920	273,920	273,920	273,920

Note: This table reports estimates of the same models as those reported in Table 3 (for mothers) for different outcomes measuring labour supply at the intensive margin. In column (1), the dependent variable is the number of working hours per week (including 0s for non-working mothers). In columns (2) to (4), the dependent variables are indicators for whether the mother works less than 16 h, between 16 and 30, and more than 30 h, respectively. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.4
Panel data estimates of the effect on mothers' labour market outcomes of the youngest child's eligibility for free childcare for different subgroups of mothers.

	(1)		(2)		(3)	
	Education		Partnership status		Local unemployment	
	main effect	* low education	main effect	* mother has partner	main effect	* low unemp in TTWA
A. Dependent variabe: Mother is in the labour force						
1st term PT	−0.007	0.023**	0.000	0.003	0.002	0.004
	(0.005)	(0.010)	(0.009)	(0.010)	(0.006)	(0.009)
2nd term PT	0.002	0.017	0.016	−0.007	0.011	−0.001
	(0.009)	(0.015)	(0.014)	(0.013)	(0.010)	(0.016)
3rd term PT	0.017	0.01	0.033**	−0.016	0.016	0.012
	(0.012)	(0.020)	(0.016)	(0.014)	(0.013)	(0.020)
4th term PT	0.016	0.015	0.036**	−0.018	0.014	0.021
	(0.014)	(0.024)	(0.018)	(0.017)	(0.015)	(0.022)
5th term PT	0.021	0.009	0.031	−0.009	0.016	0.022
	(0.017)	(0.027)	(0.021)	(0.020)	(0.017)	(0.026)
1st term FT - 3rd PT	0.035***	−0.008	0.035***	−0.005	0.028***	0.008
	(0.010)	(0.012)	(0.011)	(0.011)	(0.007)	(0.012)
2nd term FT - 3rd PT	0.060***	−0.013	0.058***	−0.008	0.041***	0.031**
	(0.013)	(0.016)	(0.015)	(0.015)	(0.009)	(0.015)
3rd term FT - 3rd PT	0.071***	−0.025	0.065***	−0.011	0.043***	0.037*
	(0.016)	(0.020)	(0.017)	(0.016)	(0.011)	(0.019)
B. Dependent variable: Mother is employed						
1st term PT	−0.008	0.009	−0.009	0.006	−0.001	−0.005
	(0.005)	(0.008)	(0.007)	(0.008)	(0.005)	(0.008)
2nd term PT	0.000	0.000	0.007	−0.010	0.003	−0.009
	(0.009)	(0.013)	(0.010)	(0.011)	(0.009)	(0.013)
3rd term PT	0.015	−0.013	0.022*	−0.019	0.013	−0.013
	(0.012)	(0.018)	(0.012)	(0.012)	(0.011)	(0.015)
4th term PT	0.02	−0.01	0.028**	−0.018	0.015	−0.001
	(0.014)	(0.022)	(0.014)	(0.015)	(0.013)	(0.017)
5th term PT	0.016	−0.004	0.026	−0.016	0.011	0.007
	(0.017)	(0.024)	(0.016)	(0.016)	(0.014)	(0.020)
1st term FT - 3rd PT	0.015*	−0.008	0.007	0.005	0.002	0.025**
	(0.009)	(0.012)	(0.009)	(0.009)	(0.006)	(0.011)
2nd term FT - 3rd PT	0.036***	−0.019	0.014	0.015	0.01	0.043***
	(0.012)	(0.016)	(0.012)	(0.011)	(0.009)	(0.015)
3rd term FT - 3rd PT	0.049***	−0.026	0.029*	0.008	0.020*	0.040**
	(0.016)	(0.019)	(0.015)	(0.013)	(0.011)	(0.018)
Observations	275,703		276,018		275,994	

Note: This table reports estimates of the same models as those reported in Table 3, where we also include interactions of all variables with subgroup indicators. In columns (1), the indicator is a dummy for whether the mother has a partner. In columns (2), the indicator is a dummy for whether the mother has low education (i.e. if her highest qualification is below A-level). In columns (3), the indicator is a dummy for whether the mother lives in a low unemployment area (if the unemployment rate in the Travel to Work Area (TTWA) in which they live is below the median unemployment rate across all TTWAs). $^{*}p < 0.10$, **p<.05, $^{***}p < 0.01$.

Table A.5
Comparison of estimates of the same parameters in the 2011 Census and in the 2010–2013 LFS.

	(1) In labour force	(2)	(3) In work	(4)
	Census	LFS	Census	LFS
A. Effects of part-time eligibility				
One term	0.004	0.003	−0.000	−0.003
	(0.008)	(0.006)	(0.008)	(0.005)
Two terms (vs one term)	−0.001	0.008	0.000	0.001*
	(0.008)	(0.007)	(0.008)	(0.006)
Three terms (vs two terms)	−0.003	0.015**	−0.005	0.007
	(0.008)	(0.007)	(0.008)	(0.005)
Four terms (vs three terms)	−0.005	−0.007	−0.002	0.001
	(0.008)	(0.008)	(0.008)	(0.006)
B. Effects of full-time eligibility relative to part-time				
2nd term (vs 4th term PT)	0.036***	0.034***	0.013*	0.015**
	(0.007)	(0.007)	(0.007)	(0.007)
Observations	183	276,018	183	276,018

Note: The estimates in the "Census" columns are copied from Table 2 for ease of comparison. The LFS coefficients are computed using estimates of a regression of a labour market outcome (indicator for labour force participation or for employment) on indicators for whether the youngest child is in a particular term of eligibility, indicators for whether any other child is in a particular term of eligibility, the number of children in the age bands 0–2; 2–4; 5–9; 10–15 in the household, age-in-month dummies of the youngest child in the household, quarter of observation dummies, whether the mother has a partner, and where we have interacted all eligibility dummies with an indicator for whether the mother is observed in the 2010–13 period. All the regressions are linear regressions with mother-level fixed effects. Based on the estimates of the model, we compute and report here estimates of the exact same parameters as those we can estimate in the Census for the 2010–13 period. Standard errors for the LFS estimates are clustered at the LEA level. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

References

Andresen, M.E., Havnes, T., 2019. Child care, parental labor supply and tax revenue. Lab. Econ. 61, 101762.
Baker, M., Gruber, J., Milligan, K., 2008. Universal child care, maternal labor supply, and family well-being. J. Polit. Economy 116, 709–745.
Bauernschuster, S., Schlotter, M., 2015. Public child care and mothers' labor supply – evidence from two quasi-experiments. J. Public Econ. 123, 1–16.
Berg, G., Bozio, A., Dias, M.C., 2020. Policy discontinuity and duration outcomes. Quant. Econ. 11, 871–916.
Berlinski, S., Galiani, S., 2007. The effect of a large expansion of pre-primary school facilities on preschool attendance and maternal employment. Lab. Econ. 14, 665–680.
Berlinski, S., Galiani, S., McEwan, P., 2011. Preschool and maternal labor market outcomes: evidence from a regression discontinuity design. Econ. Dev. Cult. Change 598, 313–344.
Berthelon, M., Kruger, D., Oyarzun, M., 2015. The Effects of Longer School Days on Mothers' Labor Force Participation. IZA Discussion Paper 9212.
Black, S., Devereux, P., Salvanes, K., 2011. Too young to leave the nest? The effects of school starting age. Rev. Econ. Stat. 93 (2), 455–467.
Blanden, J., Bono, E.D., McNally, S., Rabe, B., 2016. Universal pre-school education: the case of public funding with private provision. Econ. J. 126, 682–723.
Blanden, J., DelBono, E., Hansen, K., Rabe, B., 2021. Quantity and quality of childcare and children's educational outcomes. J. Popul. Econ. doi:10.1007/s00148-021-00835-4.
Blau, D., Hagy, A., 1998. The demand for quality in child care. J. Polit. Economy 106, 104–146.
Blundell, R., Costa-Dias, M., Meghir, C., Shaw, J., 2016. Female labour supply, human capital and welfare reform. Econometrica 84 (5), 1705–1753.
Borusyak, K., Jaravel, X., Spiess, J., 2021. Revisiting event study designs: robust and efficient estimation. https://t.co/4kDNxTdkCc?amp=1.
Brewer, M., Cattan, S., Crawford, C., 2014. State Support for Early Childhood Education and Care in England. IFS Green Budget.
Bryson, C., Brewer, M., Sibieta, L., Butt, S., 2012. The Role of Informal Childcare: A Synthesis and Critical Review of the Evidence. Nuffield Foundation report, London.
Buckles, K.S., Hungerman, D.M., 2013. Season of birth and later outcomes: old questions, new answers. Rev. Econ. Stat. 95 (3), 711–724.
Callaway, B., Sant'Anna, P.H.C., 2020. Difference-in-differences with multiple time periods. J. Econom. doi:10.1016/j.jeconom.2020.12.001.
Calonico, S., Cattanaseo, M.D., Farrell, M.H., 2020. Optimal bandwidth choice for robust bias-corrected inference in regression discontinuity designs. Econometrics J. 23, 192–210.
Cannon, J., Jacknowitcz, A., Painter, G., 2006. Is full better than half? Examining the longitudinal effects of full-day kindergarten attendance. J. Policy Anal. Manage. 25 (2), 299–321.
Cascio, E., Haider, S., Nielsen, H.S., 2015. The effectiveness of policies that promote labor force participation of women with children: a collection of national studies. Lab. Econ. 36, 64–71.
Cattan, S., 2016. Universal Preschool and Maternal Labour Supply. IZA World of Labor 312.
Clarke, D., Orrefice, S., Quintana-Domeque, C., 2019. The demand for season of birth. J. Appl. Econom. 34, 707–723.
de Chaisemartin, C., D'Hautfœuille, X., 2020. Tow-way fixed effects estimators with heterogeneous treatment effects. Am. Econ. Rev. 110 (9), 2964–2996.
Dhuey, E., Eid, J., Neill, C., 2020. Parental employment effects of switching from half-day to full-day kindergarten: evidence from Ontario's French schools. Can. Public Policy 46, 145–174.
Drange, N., Havnes, T., 2019. Early childcare and cognitive development: evidence from an assignment lottery. J. Lab. Econ. 37.
Department for Education (2013). Education provision: children under 5 years of age, January 2013, https://www.gov.uk/government/statistics/provision-for-children-under-5-years-of-age-in-england-january-2013, Main Tables SFR23/2013.
DWP ONS, Social and Vital Statistics Division, NatCen Social Research, 2016. Family Resources Survey, 2005/06-2013/14 and Households Below Average Income, 1994/95-2013/14: Safe Room Access. [data collection]. 5th edition. UK data service. SN: 7196. 10.5255/UKDA-SN-7196-6
Felfe, C., Lechner, M., Thiemann, P., 2016. After-school care and parents' labor supply. Lab. Econ. 42, 64–75.
Fitzpatrick, M., 2010. Preschoolers enrolled and mothers at work? The effects of universal prekindergarten. J. Lab. Econ. 28, 51–84.
Gathmann, C., Sass, B., 2018. Taxing childcare: effects on childcare choices, family labor supply, and children. J. Labor Econ. 36, 665–709.
Gelbach, J.B., 2002. Public schooling for young children and maternal labour supply. Am. Econ. Rev. 92, 307–322.
Goodman-Bacon, A., 2021. Difference-in-differences with variation in treatment timing. J. Econom. doi:10.1016/j.jeconom.2021.03.014.
Goux, D., Maurin, E., 2010. Public school availability for two-year olds and mothers' labour supply. Lab. Econ. 17, 951–962.
Havnes, T., Mogstad, M., 2011. Money for nothing? Universal child care and maternal employment. J. Public Econ. 95, 1455–1465.
Lefebvre, P., Merrigan, P., Verstraete, M., 2009. Dynamic labour supply effects of child care subsidies: evidence from a Canadian natural experiment on low-fee universal child care. Lab. Econ. 16, 490–502.
Lundin, D., Mork, E., Ockert, B., 2008. How far can reduced child care prices push female labour supply? Lab. Econ. 15, 647–659.
Mumford, K., Parera-Nicolaus, A., Pena-Boquete, Y., 2020. Labour supply and childcare: allowing both parents to choose. Oxford Bull. Econ. Stat. 82, 577–602.
Nollenberger, N., Rodriguez-Planas, N., 2015. Full-time universal childcare in a context of low maternal employment: modest but persistent effects. Lab. Econ. 136, 124–136.
Dhuey, E., Lamontagne, J., and T. Zhang, T. (2019). The Impact of Full-Day Kindergarten on Maternal Labour Supply. IZA DP 1250.
Powell, L., 2002. Joint labor supply and childcare choice decisions of married mothers. J. Hum. Resour. 37, 106–128.
Shure, N., 2019. School hours and maternal labor supply. Kyklos 72 (1), 118–151.
Yamaguchi, S., Asai, Y., Kambayashi, R., 2018. Effects of subsidized childcare on mothers labor supply under a rationing mechanism. Lab. Econ. 55, 1–17.