INSTITUTE OF ECONOMICS
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Közgazdaság- és Regionális Tudományi Kutatóközpont
KÖZGAZDASÁG-TUDOMÁNYI INTÉZET

KRTK-KTI WORKING PAPERS | KRTK-KTI MŰHELYTANULMÁNYOK

The impact of childcare on maternal employment

BENCE SZABÓ – JUDIT BEREI – MÁRTON CSILLAG –

HANNA ERŐS – JUDIT KREKÓ –ÁGOTA SCHARLE

KRTK-KTI WP – 2022/20

October 2022

https://kti.krtk.hu/wp-content/uploads/2022/10/CERSIEWP202220.pdf

INSTITUTE OF ECONOMICS, CENTRE FOR ECONOMIC AND REGIONAL STUDIES
EÖTVÖS LORÁND RESEARCH NETWORK (ELKH)
BUDAPEST, 2022

ABSTRACT
This paper examines the effect of childcare availability on maternal employment in Hungary based on 2016 Microcensus data. We exploit the exogenous variation in access to childcare due to informal admission practices based on the date of birth, to identify the effect of childcare availability on maternal employment and the children's enrolment. We find that on average, expanding the coverage of nurseries to the same level as kindergartens would lead to around 7.3 percentage points higher maternal employment, an around 25% higher employment rate compared to the baseline of mothers with a child aged 2-2.5 years. At the same time, the decomposition of the link between childcare availability and employment shows that enrolment would increase by 17.7 percentage points due to the higher coverage, close to 40% compared to the baseline. Enrolment in childcare would increase maternal employment probability by around 41 percentage points, around two-thirds of the employment rate of mothers. We also examine the heterogeneities of the effect along demographic characteristics using causal forests, and the economic cycle by expanding the analysis to the 2011 Census. We find that in 2016 the childcare availability effect is higher for mothers with 3 children, living in villages, or municipalities without nurseries. The employment effect is lower in the 2011 Census, while the effect on enrolment in formal childcare remains similar, suggesting the importance of weaker labour demand in 2011.

JEL codes: J13, J22
Keywords: mothers' labour supply, childcare availability

BENCE SZABÓ
Bocconi University
e-mail: benceszabo@phd.unibocconi.it

HANNA ERŐS
Budapest Institute for Policy Analysis
e-mail: hanna.eros@budapestinstitute.eu

JUDIT BEREI
Budapest Institute for Policy Analysis
e-mail: judit.berei@budapestinstitute.eu

JUDIT KREKÓ
KRTK KTI and Budapest Institute for Policy Analysis
e-mail: kreko.judit@krtk.hu

MÁRTON CSILLAG
KRTK KTI and Budapest Institute for Policy Analysis
e-mail: csillag.marton@krtk.hu

ÁGOTA SCHARLE
Budapest Institute for Policy Analysis
email: agota.scharle@budapestinstitute.eu

A bölcsődei ellátás hatása az anyák munkavállalására

SZABÓ BENCE – BEREI JUDIT – CSILLAG MÁRTON –

ERŐS HANNA – KREKÓ JUDIT – SCHARLE ÁGOTA

ÖSSZEFOGLALÓ

A tanulmány a bölcsőde vagy óvoda elérhetőségének hatását vizsgálja az anyák foglalkoztatására Magyarországon, a 2016-os Mikrocenzus adatai alapján. A bölcsődéknél jóval szélesebb körben elérhető óvodák informális felvételi szabályait kihasználva, oksági becslést adunk arra, hogy a három év alatti gyermekek intézménybe járásának a lehetősége hogyan hat a gyermekek tényleges beíratására és az anyák foglalkoztatási arányára. Eredményeink szerint a bölcsődei férőhelyek megnövelése az óvodához hasonló szintre mintegy 7,3 százalékponttal növelné az érintett anyák foglalkoztatottságát, ami a 2-2,5 éves gyermeket nevelő anyák foglalkoztatási rátájához viszonyítva mintegy 25%-kos emelkedést jelent. Ugyanakkor a napközbeni ellátás férőhelyszáma és a foglalkoztatás közötti kapcsolat dekompozíciója azt mutatja, hogy az intézmények kihasználtsága a férőhelyek bővülése esetén 17,7 százalékponttal, közel 40%-kal növekedne az alaphelyzethez képest. A gyermekek intézménybe járása pedig mintegy 41 százalékponttal, kétharmaddal növelné az anyák foglalkoztatását. A hatás demográfiai jellemzők mentén mutatkozó heterogenitását is vizsgáljuk az ok-okozati erdők (random forest), valamint a gazdasági ciklus mentén, kiterjesztve az elemzést a 2011-es népszámlálásra. Azt találjuk, hogy 2016-ban a napközbeni ellátás elérhetőségének hatása magasabb a 3 gyermekes, a falvakban élő és a bölcsőde nélküli településeken élő anyák esetében. A 2011-es népszámlálási adatokon a foglalkoztatásra becsült hatás alacsonyabb, míg az óvodai beiratkozásra gyakorolt hatás hasonló, ami a 2011-es gyengébb munkaerő-kereslet jelentőségét tükrözheti.

JEL: J13, J22

Kulcsszavak: anyák munkakínálata, bölcsődei és óvodai ellátás

1. Introduction[1]

The employment rate of mothers with children aged below 3 has been very low in Hungary, around a third of the OECD average, while the share of those working is much higher (nearing the OECD average) for mothers with children aged 3 to 5. (OECD, 2020). This may partly be explained by the parental leave system, which provides paid leave at a relatively high replacement rate for 24 months and an additional period of 12 months paid at a low flat rate. However, the large differential in nursery and kindergarten coverage rates suggests that the availability of childcare for children aged below 3 also contributes to explaining the employment gap among mothers. Indeed, based on data from 1998-2011, Lovász and Szabó-Morvai (2019) find that access to subsidized childcare could raise the maternal labour supply by 24 per cent in Hungary. While childcare availability has improved substantially between 2005-2019, the Barcelona targets have not been reached. Furthermore, there are large inequalities in access to childcare for children under 3 years of age: availability is close to 30% in cities, and below 10% in rural areas.

Long absences from work can result in a depreciation of human capital and can have detrimental effects on women's subsequent careers (Blundell et al, 2016, Bálint - Köllő 2008). Maternal employment tends to reduce child poverty (Dotti Sani and Scherer 2018, Thévenon et al. 2018) and also supports female emancipation (Korpi, Ferrarini, and Englund 2013). Raising mothers' employment rate has also become a pressing economic issue in Hungary as labour shortages intensified before the COVID-19 pandemic, which motivated the government to expand childcare services and increase financial incentives for women's return to work.[2]
In this paper, we investigate the effect of childcare availability on maternal employment, building on the work of Lovász and Szabó-Morvai (2019). This approach exploits the fact that kindergartens are substantially more available than nurseries, which induces an age-specific discontinuity in access to childcare. The timing of kindergarten admission (i.e. access to formal childcare), depends on the birth date of the child, which can be regarded as random. Specifically, we use the fact that, due to national guidelines and a wide-spread informal rule, children born between September and December are much more likely to be admitted to kindergarten before they turn 3 years of age than kids born between January and August. This approach helps to overcome the potential endogeneity of regional differences in childcare capacity and enables us to separate the effect of better access to childcare from the loss of paid leave (and possibly of social expectations), occurring when the child turns three.

The paper contributes to existing research in three areas. First, we estimate the effect of childcare on mothers' employment over a period characterised by changes in labour demand as well as in the policy context (including a sizeable expansion of childcare availability, a reduction in the age of obligatory kindergarten enrolment, and a gradual relaxation of rules prohibiting paid work while on parental leave). These changes have altered incentives for mothers' return to work, calling for an update of the earlier analysis by Szabó and Morvai (2019). Second, as we have data from both 2016 and 2011, it is possible to discern the potential difference in the effect of childcare expansion across a deep recession and a period of rapid recovery. Third, we can refine existing results by using a dataset that records actual childcare attendance. This allows us to investigate the mechanism through which childcare availability impacts mothers' employment and separate the likelihood of taking up child-care from the likelihood of subsequent return to work. Last, given that we have access to a large sample, we can estimate

[1] This paper was prepared using datasets from the Census, Microcensus and the Labour Force Survey of the Central Statistical Office. We thank the Central Statistical Office for providing access and the Databank of the Centre for Economic and Regional Studies for help with the administrative database and using the data room. We thank for Anna Lovász and Ágnes Szabó-Morvai for sharing the codes of their earlier analysis. Some parts of the analysis was prepared for a report commissioned by the OECD. Csillag and Krekó were supported by the „Lendület" program of the Hungarian Academy of Sciences (grant number: LP2018-2/2018).

[2] Specifically, from 2014 onwards, mothers (parents) are allowed to return to work after the child has turned 6 months old, and still keep all of the childcare benefits, which is relatively generous until age of 2 of the youngest child.

heterogeneities across different groups of mothers. This allows us to understand which groups of women are the most willing to return to work (and use formal childcare).

We find that in our main specification childcare coverage in nurseries similar to kindergartens (around 100 percentage points increase) would lead to around 7.3 percentage points higher maternal employment, an around 25% higher employment rate compared to the baseline of mothers with a child 2-2.5 years of age. At the same time, the decomposition of the link between childcare availability and employment shows childcare facility attendance would increase by 17.7 percentage points due to the higher coverage, close to 40% compared to the baseline. The ability to place the child in childcare would increase maternal employment probability by around 41 percentage points, around two-thirds of the employment rate of mothers (with a child of age 2-3). We also examine the heterogeneities of the effect along demographic characteristics using causal forests, and economic circumstances by expanding the analysis to the 2011 Census and the 2005-2020 waves of the Labour Force Survey (LFS). We find that in 2016 the childcare availability effect is higher for mothers with 3 children, living in villages, or municipalities without nurseries (the latter two largely overlap). While for employment the LFS estimates are close to the 2016 Microcensus, the effect size is lower in the 2011 Census, with the effect on childcare facility attendance remaining similar, suggesting the importance of weaker labour demand in 2011.

The remainder of the paper is structured as follows. Section 2 reviews the relevant literature. Section 3 overviews the institutional background, and Section 4 and 5 describes the Data and developments of childcare capacity. Section 6 presents the empirical strategy and Section 7 summarises the results and robustness checks. Section 8 concludes.

2. Existing research on the impact of childcare on mothers' labour market outcomes

There is a large empirical literature on the relationship between childcare availability and mothers' labour market outcomes, with estimates from specific countries varying greatly. For Hungary, Lovász and Szabó-Morvai (2019) find that increasing the availability of nurseries to the same (high) level as kindergartens would increase maternal employment rates by 24 per cent. Kunze and Liu (2019) investigate the impact of a reform expanding childcare by 10 percentage points for 1-2-year-olds in Norway using a reduced-form estimation approach and find that it led to a short-term increase of 3.4% and a long-term increase of 2.8% in maternal employment for the targeted group. A study from Italy on childcare below three by Carta and Rizzica (2018) exploits discontinuities in the eligibility rules and the staggered implementation of the reform to analyse the effect of an expansion of access to highly subsidized child care for 2-year-olds and reports an increase of 5–7 percentage points on the labour participation of mothers. Effect sizes of policy reforms for children three and above tend to be smaller (Nollenberger and Rodríguez-Planas 2015; Givord and Marbot 2015). Moreover, most of the above-mentioned studies find greater effects for mothers with two or more children, who are likely even more time-constrained and/or may have completed fertility and might therefore respond more to a reform (Kunze and Liu 2019; Nollenberger and Rodríguez-Planas 2015; Givord and Marbot 2015).

Several key reasons have been identified in the literature to explain the diverging estimates. One of these factors is whether a policy has a high or low potential impact to begin with, depending on the initial level of maternal labour supply. Lovász and Szabó-Morvai (2019) review the literature from this aspect. On the one hand, they find that studies on low potential countries (i.e. high initial maternal employment) identify smaller or no effects of increasing childcare availability on maternal labour supply (Fitzpatrick 2010; Givord and Marbot 2015). On the other hand, in countries where the pre-existing level of maternal employment is low, and mothers are likely to be constrained by the low availability of early childcare, studies find a higher impact of expansion policies – including in Hungary (Nollenberger and Rodríguez-Planas 2015; Lovász

and Szabó-Morvai 2019). Furthermore, the cross-country analysis of Szabó-Morvai and Lovász (2017) suggests that the effect of childcare facilities is greatest in Central Eastern European countries where the pre-existing level of maternal employment is low, and maternal leave ends around age 3 of the child, while the effects are smaller in Southern and Western EU countries with higher employment levels, and shorter maternal leave.

Besides initial maternal labour supply, relevant factors affecting the results include the targeted age group (children below three or children between three and school-age), the institutional context of the country, the quality and cost of childcare, as well as social norms dictating the length of time a mother should spend home with a child. Cattan (2016) and Vuri (2016) both discuss how these factors interact in an international context and find that the effectiveness of childcare expansion policies is largely determined by the country-specific combination of pre-reform childcare arrangements, maternity leave regulations, and gender preferences. More specifically, a childcare reform tends to be more effective in an environment with high-quality childcare facilities, fewer options for informal childcare arrangements, shorter maternity leave regulations, and a more egalitarian gender view, while it tends to be less effective in countries with the opposite characteristics, including more traditional gender roles (Vuri 2016; Cattan 2016). According to Kleven et al. (2019; 2020), social and gender norms play a profound role in maternal employment. Investigating the joint impact of an enormous expansion in childcare subsidies and parental leave in Austria since the 1950s, they find that the state intervention policies had effectively no impact on gender convergence in earnings (Kleven et al. 2020). Similarly, they discover strikingly high correlations between prevailing traditional gender views (high fraction of people agree that women with children under school age or in school should stay at home) and long-run child penalties in earnings when estimating the effect across countries using event studies around the birth of the first child (Kleven et al. 2019). This aspect is especially important in Hungary where traditional gender norms still largely determine how long a mother should stay at home after birth.

This paper studies both the effect of childcare expansion on employment and its effect on nursery or kindergarten attendance. The latter, attendance has been less examined in the literature, although it provides useful compliance information, that is, how many children are indeed taken to kindergarten if those places are unofficially available. One of the few exceptions is Bousselin (2022) who finds that for a 1% increase in day care attendance after a reform in Luxembourg, there is a 0.17 percentage point increase in the employment rate of mothers with very young children. Another paper from Argentina by Berlinski and Galiani (2007) uses a differences-in-differences strategy to identify the impact of a large early childcare expansion on attendance and maternal labour supply and reports sizeable effects on participation among children aged 3 and 5 as well as an increase in employment. Furthermore, Bauernschuster and Schlotter (2015) use a very similar identification strategy as the current paper to study the introduction of a legal claim to a place in a kindergarten in Germany. With the help of an instrumental variable approach that is enabled by day-of-birth cut-off dates, they show that eligibility for public child care increases mothers' labour supply by 6 percentage points whereas if a youngest child actually attends kindergarten (as a result of this cut-off), the mother's probability of being employed increases by around 35 percentage points (Bauernschuster and Schlotter, 2015).

Moreover, the current paper also tests for heterogeneity in mothers' characteristics by attendance. While other studies also investigate heterogeneous effects on the employment outcome of mothers, our findings are unique in that they additionally discover which mothers place their children in early childcare. There is a substantial difference in interpretation between these two; there might be mothers who place their children in kindergarten as a consequence of the expansion but are unwilling or unable to find employment. Both effects are important in understanding the implications of any policy reform.

When looking at heterogeneous effects among employed mothers, studies tend to find diverging effects for different subgroups. Bauernschuster and Schlotter (2015) find that the expansion

effects are largest for mothers with above median years of schooling, mothers with above median age, mothers also having older children as well as mothers whose youngest child's age distance to his oldest sibling is above the median. Carta and Rizzica (2018) – who study a reform that introduced early access to subsidized childcare for 2-year-old children in Italy – discovers significantly larger effects among more affluent families and among married women (who, traditionally, have a lower labour market participation rate). Moreover, most studies agree that the employment effect is strongest among women with more than one child but only when the mother has no younger children than the one who is eligible for kindergarten after the reform (Bauernschuster and Schlotter 2015; Nollenberger and Rodríguez-Planas 2015; Carta and Rizzica 2018; Kunze and Liu 2019).

3. Institutional background[3]

Daycare for pre-school children is traditionally highly regulated and mainly provided by municipalities in Hungary (OECD 2004, OECD 2006, Eurydice 2022). Daycare is divided into two distinct institutional forms: nurseries for children aged below three and kindergartens for children aged between three and starting school (around age 6 or 7). While access and enrolment in kindergartens have been above 90% since the mid-1990s, access to nurseries remained below 10 per cent (well below the Barcelona target of 33%) until around 2012.

Nurseries are predominantly run by local governments and are located in urban centres. The number of nursery places dropped sharply during the transition to a market economy in the early 1990s and stagnated until 2006 when it started to increase slowly.

The provision of childcare for children aged below three was not a priority of public policy until recently, as there was cross-party consensus about keeping the long parental leave (available until the child turns three), and, given that unemployment remained high for decades after the transition, there seemed to be no need to increase maternal labour supply.

Nevertheless, there were piecemeal reforms, largely initiated by professionals, to support the development of nurseries from the mid-1990s (see further details in the Appendix). Hungary's accession to the EU in 2004 gave further impetus to this trend as the EU encouraged and provided funding for social investment in childcare.

Nurseries were fully financed by municipalities until 1997 when a per capita subsidy was introduced. In the same year, a new, more flexible institution was introduced, the so-called *family daycare centre*, which was intended to improve access to daycare in rural areas (see Appendix for more detail). This format (though tightly regulated) was made eligible for a per capita subsidy from the central government in 2003. Capacity in such family daycare centres increased steadily until 2017, but their share in enrolment of nursery-age children remained below 20%.

Starting around 2006, successive governments launched competitive grant-schemes to encourage local governments to establish nurseries (mainly using EU funds). These programmes gained impetus, especially after 2012 when economic growth accelerated and the shortage of skilled labour became more pressing. The introduction of parental fees in nurseries in 2012, though highly regulated and relatively small, was also intended to generate funding to enable municipalities to maintain daycare capacities. This was further strengthened by a new rule enacted in 2017 (enforced from 2020), which obliged municipalities to regularly assess demand and provide daycare if there are at least 40 children of nursery age living in the municipality (or at least 5 parents request the service) (see Appendix).

[3] See the Appendix for an overview of child-related cash benefits.

There were also several measures to ease the flexible use of the existing infrastructure. First, the professional guidelines for maintaining nurseries issued in 1998, allowed nurseries to exceed the legal maximum for the number of children per group by 20%, so that the effective maximum group size was 12 (rather than 10). This practice was formalised in 2010 when the legally prescribed maximum was raised to 12 (or even 14 if all children are aged 2), and the updated guideline issued in 2011 stated that no deviation is acceptable.[4] To facilitate the flexible use of excess capacity in kindergartens, a new institutional form was introduced in September 2009, which allowed kindergartens to set up mixed-age groups in which they could enrol a maximum of 5 children aged between 2 and 3 (along with 15 children aged 3 or over). This format was dismantled in 2017. As of September 2011, kindergartens were allowed to enrol children aged 2.5 without any restrictions on their share within a group.[5] It should be noted that excess capacities were curbed by a later rule that lowered the compulsory enrolment age in kindergarten from 5 to 3, as of September 2015.

In 2017, the government revised the regulation of nurseries and introduced new forms of service provision. All new forms were labelled "nursery" but service content was quite varied:

(1) traditional nurseries were left largely unchanged
(2) mini-nurseries were similar to traditional nurseries except that the maximum group size is 7 (or 8 if all children are aged 2)
(3) family-nurseries were supposed to replace family day-care centres, under similar conditions except that they could only take children aged up to three (while previously the upper limit was age 14 for them)
(4) employer's nurseries that are similar to family-nurseries regarding group size, the required infrastructure and staff qualifications.

The supply of flexible forms of daycare declined slightly after 2017, most likely because the earlier formats were dismantled while the establishment of new institutions required considerable administration (while traditional nurseries were left unchanged). In the rest of the paper, all data concerning nurseries include both nurseries and family day care centres (before 2017) or the above described four institutional forms (as of 2017).

4. Data sources

To estimate the effect of childcare availability on maternal employment, we use the 2016 Microcensus of Hungary[6], augmented by the 2011 Census of Hungary and the T-STAR Hungarian regional data, provided by the Databank of the Centre for Economic and Regional Studies. The 2016 Microcensus contains a 10% anonymized sample of the population, and is representative at the micro-region (LAU2) level units level according to age groups, education, employment, marital status and inhabited residences, while the 2011 Census assures full coverage. The T_STAR is compiled based on administrative reports by the Central Statistical Office of Hungary, we extract information on nursery and kindergarten places, as well as the number of children in each locality[7].

Primarily, we identify the youngest children of the family units in the 2016 Microcensus, and we calculate several parental attributes such as age, education, labour force status or place of living. Our key outcome is the employment rate of mothers: Please note that this is self-reported employment status, it is not equivalent to the ILO employment criterion. As an additional

[4] Note that an interim guideline issued in 2009 prescribed the acceptable maximum to be 12-14 children per group.
[5] Both provisions applied to excess capacity only, which was enforced by elaborate rules to ensure that children aged 3 were given priority in kindergartens (see more detail in the Appendix).
[6] The 2016 Microcensus contains a 10% anonymized sample of the population, representative at the micro-region (LAU2) level according to age groups, education, employment, marital status, and inhabited residences.
[7] Please note that this is the number of permanently registered persons, which is not necessarily equivalent to the number of children actually living in the given settlement.

outcome, we also look at enrolment in formal childcare of the child, to form an idea about the main channel via which institutional capacity could affect the labour market status of mothers.

The main advantage of the Microcensus compared to the Labour force Survey, on which the estimations of Lovász and Szabó-Morvai (2019) are based, is that we observe actual childcare attendance of the children, hence we can unravel the mechanism behind the impact of childcare availability. Nevertheless, as a robustness check, we also estimate our main specification on datasets of the 2011 Census, and several waves of the Labour Force Survey (See section 7.4).

5. Evolution of childcare capacities and labour market for mothers

In what follows, we show the evolution of childcare capacities based on nursery catchment areas. These catchment areas were developed based on data from the 2011 Census, and are fixed for the entire period discussed in this paper (2005-2019)[8].

We measure childcare availability by local nursery[9] and kindergarten coverage relevant to the child's place of living. First, we construct nursery and kindergarten catchment areas based on the 2011 Census, which is indicated at the municipality level where children live, and where they attend nursery or kindergarten. Each municipality with a nursery is the centre of its area, and then based on the commuting information in the 2011 Census we assign each municipality to a unique centre, the one where most children from the municipality would attend a nursery.[10] In practice, this results in 493 areas mostly centred around a locally important town or city.[11] Then using the T-STAR regional data for each of these areas we calculate childcare availability as the proportion of nursery places to the number of 0–2-year-olds, and the proportion of kindergarten places to the number of 3–5-year-olds for 2016. Figure 1 displays these catchment areas along with the nursery coverage in the area for the year 2011.

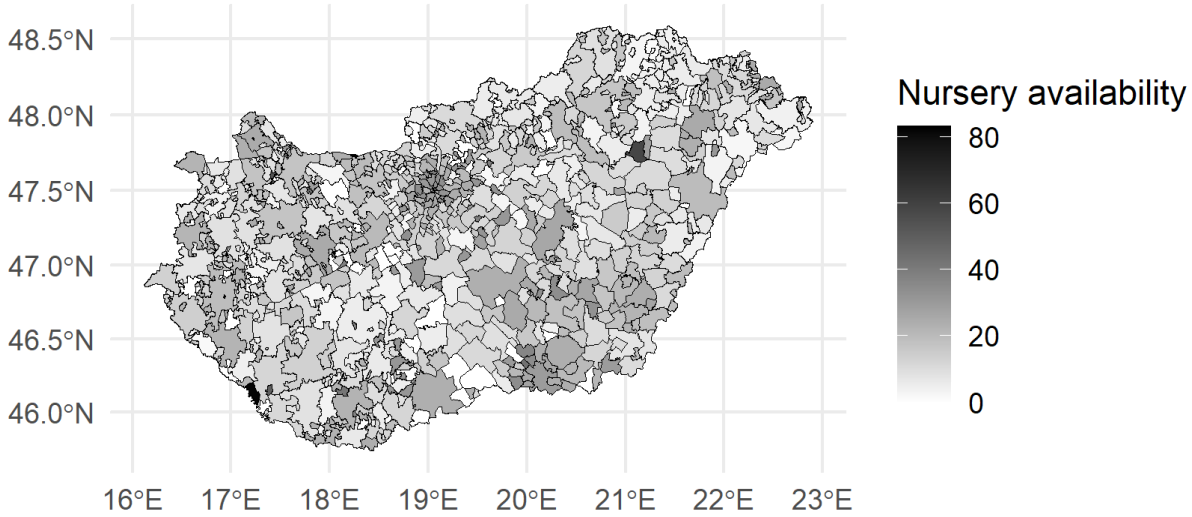
[8] Note that by 2011 nursey capacities have developed significantly, so it is possible that prior to 2010, (a) there are zero places available in a nursery in a given catchment area, and (b) we would have found fewer catchment areas empirically.

[9] Pls note that we also include family day care centres when calculating nursery places. While these centres were not exclusive to children below age 3, the vast majority of those attending these centres were of nursery age.

[10] For municipalities without children in the data, we assign the closest centre.

[11] Pls notice that this is far fewer than the number of municipalities (LAU2, 3152 units), meaning that there was one nursery for (roughly) every six municipalities. Catchment areas a much more numerous than the number of micro-regions (LAU1, 174+23 units), which are typically considered a local labour market and are the basis of local governance and the distribution of EU Funds. Hence, we will be able to control for micro-region fixed effects in out analysis.

Figure 1: Nursery catchment areas in Hungary based on commuting information from the 2011 Census, and nursery coverage rate in %



Notes: The authors' calculations are based on the T-STAR regional data on municipal information of nursery places and population, and the 2011 Census of Hungary.

We briefly present the evolution of nursery coverage rates (as well as the difference between nursery and kindergarten coverage rates) in Table 1 below. Before looking at its distribution, we present the evolution of nursery and kindergarten places, as well as the number of children in the two relevant age categories. Please note that we adjust nursery places available before 2010 to take into account the change in regulation regarding overbooking, hence, we multiplied places available by a factor of 1.2 before 2011, and this was exclusive to state-run nurseries.

Table 1: Number of children and childcare capacity

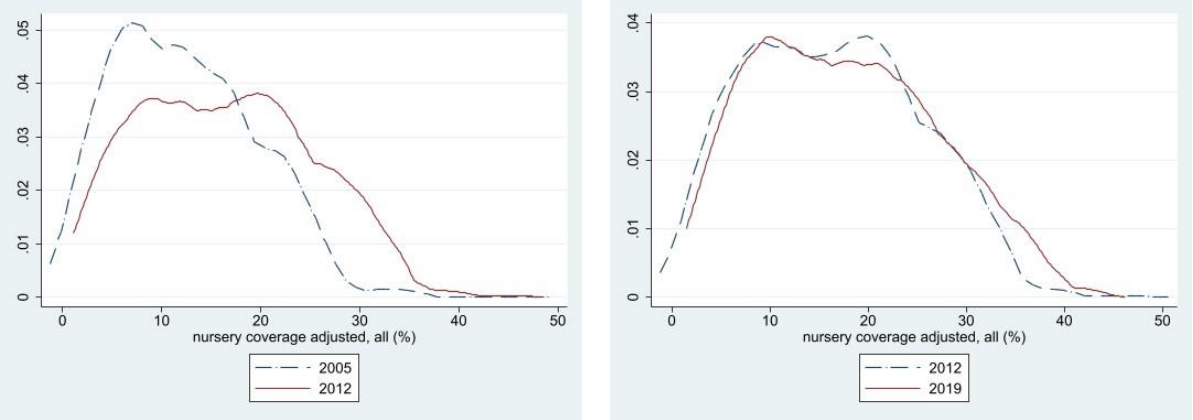
year	number of children aged 0-2	number of children aged 3-5	number of available nursery places	number of available kindergarten places
2005	284977	289846	28624	349564
2006	290865	287507	29106	351710
2007	293258	286058	29940	349404
2008	295040	287465	32650	354157
2009	291749	293107	34786	362854
2010	284462	295777	37387	369966
2011	273731	297660	41715	374717
2012	267952	294589	44070	377001
2013	267832	287971	45713	378402
2014	273608	277474	46835	380288
2015	277029	272411	48027	379456
2016	283072	272811	48069	379132
2017	284179	279142	46475	379124
2018	283145	282874	47162	382491
2019	279802	288713	48688	386011

Notes: Aggregate figures are based on T-Star. The number of nursery places includes both nurseries and family day care centres. In years before 2011, the number of subsidized nursery places is adjusted by a factor of 1.2 to take into account potential overbooking.

First, we can see a marked increase in the number of nursery places available between 2007 and 2012, with an almost 50 per cent growth over a 5-year-period; and a further slower expansion between 2012 and 2019. It is worth noting that about half of the increase in nursery capacities can be ascribed to family day care centres. Second, there was a slow gradual increase in kindergarten capacities throughout the 15 years examined. Third, the number of children reflects the dynamics of childbearing, in particular the marked fall in births in the aftermath of the great recession. We can see that for the 0-2 age group, this reaches its minimum in around 2012-2013; while for age 3-5, this is (naturally) in 2015-2016. It is thus important to take into account that coverage rate dynamics are partly a result of these demographic phenomena.

In Figure 2, we show the kernel density of nursey coverage rates (weighted by the number of children aged 0-2) for selected years: 2006, 2011 and 2017. We can see that by 2011, the coverage rate for a large number of catchment areas was in the region of 20 per cent, while this meant a very high value five years earlier. In other words, while in 2006 only one in ten children lived in a catchment area with a coverage rate above 20 per cent, by 2011, this changed to one in four children. Furthermore, in a non-negligible number of catchment areas (where 5% of young children lived), the coverage rate was 30 per cent or above. We see further small improvements by 2017, with the median coverage rate growing to 17% (from around 14%). There was no substantial improvement in the access to early childhood care in the lower tail of the distribution, with one in ten children still living in catchment areas with coverage rates of 5 per cent or below. By contrast, there were one in ten children living in catchment areas with coverage rates above 30 per cent, growing from only 5% of children living in such an area.

Figure 2: Distribution of nursery coverage rates, 2005, 2012, 2019 (%)



Source: own calculations based on T-Star

In Table 2, we show nursery coverage rates broken down by the size of the settlement (the size of the catchment area centre).[12] First of all, we can notice that early childcare supply varies widely across urban and more rural areas, and that coverage rates were much higher in the largest cities (above 40 thousand inhabitants). This meant rather extreme inequalities: only in towns with a population above 10 thousand inhabitants was the coverage rate above 5 per cent; and the difference between the most rural catchment area centres (with 2,000 inhabitants at most) and the largest cities was ten-fold (or 15 percentage points).

[12] It is worth emphasizing that coverage rates are partly determined by the number of births (and the distribution thereof). Indeed, part of the reason why coverage rates increased between 2008 and 2012 is due to the fall in the number of children age 0-2 (which in tun was a result in the decrease in the number of birth in the aftermath of the great recession), from 295 to 268 thousand children (an almost 10% decrease). Similarly, part of the stagnation of coverage rates after 2012 is a result of the increase in the number of young children to 284 thousand in 2017.

Table 2: Availability of childcare in nurseries by settlement population size

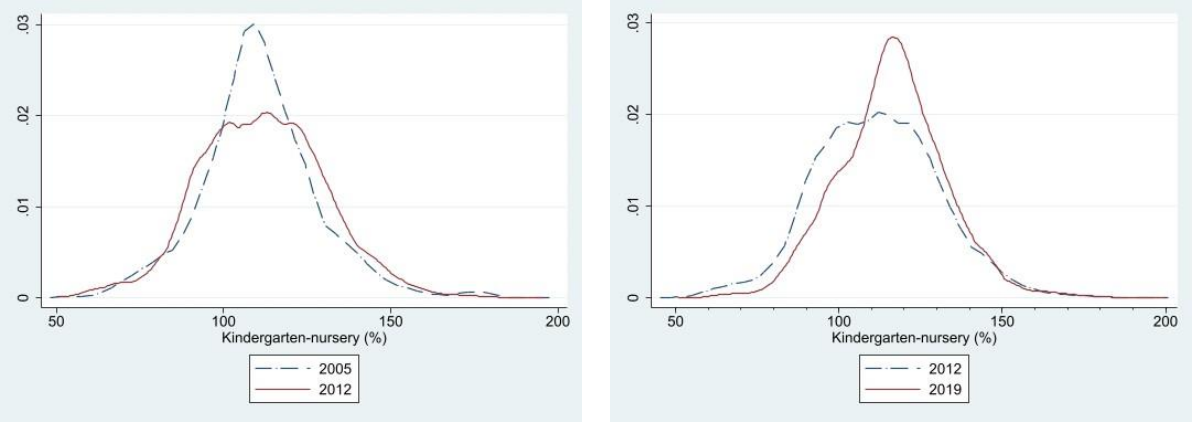
	Population size					
year	40,001+	40,000-10,000	10,000-5,000	5,000-2,000	2,000-1	Total
2005	17,4	6,8	3,6	2,9	1,2	10,0
2006	16,8	6,9	3,7	3,0	1,6	10,0
2007	16,7	7,2	3,9	3,2	2,1	10,2
2008	17,1	8,1	4,9	5,1	4,4	11,1
2009	18,0	8,7	5,5	6,3	6,3	11,9
2010	19,4	9,4	6,5	7,9	10,0	13,1
2011	21,5	11,2	8,2	10,3	12,8	15,2
2012	22,9	12,7	8,7	11,2	12,8	16,4
2013	24,0	13,4	9,0	10,8	10,9	17,1
2014	24,2	13,6	9,1	10,7	10,0	17,1
2015	24,6	13,9	9,3	10,5	9,1	17,3
2016	24,5	13,4	8,9	10,4	8,5	17,0
2017	23,9	12,8	8,4	9,8	7,5	16,4
2018	24,5	13,1	8,6	10,1	8,3	16,7
2019	25,4	13,7	9,4	10,9	9,9	17,4

Notes: Calculations are based on the T-Star database. Catchment areas are characterised by the permanent population of the centre of the nursery catchment area.

Second, the increase in capacities had some levelling effect, with the coverage rates in smaller (more rural, especially in settlements with 5000 or fewer inhabitants) settlements growing significantly quicker than in towns, and as a result, nursery coverage rates were on average around 10 per cent in all catchment area types. Third, we can see that from around 2012, there was no significant increase in overall nursery coverage rates, and there was no further decrease in coverage inequality across larger and smaller settlements. Only in towns with above 10 thousand inhabitants was there a small increase in coverage rates.

We finally turn to the difference between coverage rates in kindergartens and nurseries. In Figure 3 we display kernel densities of this difference for the same years as above. This Figure shows that between 2006 and 2011, there was a decrease in the difference between nursery and kindergarten coverage rates, and a significant increase in the variability of this difference (especially in the 'middle' of the distribution). However, in the later years, the difference between nursery and kindergarten coverage rates widened again (with the average difference increasing from 111 percentage points to 119 percentage points). This shows that indeed, the potential for children below age 3 to attend kindergarten largely expanded the supply of formal childcare availability for this specific age group.

Figure 3: Distribution of the difference between kindergarten and nursery coverage rates, 2005, 2012, 2019 (%)



Source: own calculations based on T-Star

6. Empirical strategy

We aim to provide a causal estimation of the impact of childcare availability on maternal labour supply. The access to childcare in a given region or settlement might be endogenous, as the authorities are more likely to expand the capacity of childcare institutions in regions with better labour market prospects or stronger demand. Thus, we apply an instrumental variable estimation strategy based on Lovász and Szabó-Morvai (2019) to address the endogeneity of childcare capacity. Their estimation exploits that the coverage ratio of nurseries and kindergartens sharply differ: while the total capacity of nurseries is around 17% of the 0-3-year-old population is 2016, the average coverage ratio of kindergartens exceeds 130% (see Section 4). Hence those children who are eligible to attend kindergarten have a much higher chance to access any childcare institution than children who are still eligible to nurseries. The exact date of admission into kindergarten (thus the availability of formal childcare), depends on the birthday of the child, which can be regarded as random. Consequently, following Lovász and Szabó-Morvai (2019) we use the month of birth of the child as an instrument for eligibility for high-capacity kindergartens.

In the Census and Microcensus, we observe mothers and their children on one single date of the year, the 1st of October, just after the beginning of the school year. For this observation date, there are three potential cutoff dates to consider. [13]

To support the identification strategy, let us consider Figure 4 showing the institutional attendance rate of children by age in months, as of October 1st, 2016, along which we display different cut-off points for possible admittance. The *end-August* cut-off point represents the legally mandated age threshold. In the year when children celebrate their third birthday by 31 August, they are mandated to start kindergarten education (and the kindergarten is obliged to admit her or him) in September, the beginning of the school year since 2015. A literal interpretation of the law implies that children who turn three years old only in September must be admitted into kindergarten only in September of next year. However, it is clear from the figure that even for younger children kindergarten attendance is quite high, which suggests that other rules are more relevant concerning parental and institutional behaviour.

Two additional cut-off points might govern admission decisions: one at *the end of February*, and one at *the end of December*. According to the law, introduced in 2010, children who have

[13] The definition of the controls and treatment groups differs from Lovasz and Szabó Morvai (2019) partly because of significant changes in the admission rules and because of the differences between the structure and time span LFS ans Census. See the details in Appendix A.5.

reached the age of 2.5 years can also be admitted to a kindergarten, but this is subject to consideration by the kindergarten and depends on the available excess capacity. It implies that children who turn three by the end of February might be admitted to kindergarten in September of the previous year. However, the end-December informal cut-off seems to be even more important in enrolment decisions: in practice, kindergartens usually admit children in September who turn three by the end of that calendar year. We can see that at the December cut-off nursery attendance decreases from a steady level of around 30%, while kindergarten attendance increases more than the previous trend would imply. Proceeding with the December cut-off point and using a 3-month time window around it, we could inspect the change in maternal employment in the 2016 Microcensus data (Figure 5). While before the cut-off (meaning for youngest children) the employment rate of mothers fluctuates around 30%, after the cut-off it elevates to a level of around 45%, driven by the mothers of children born in November and October. We assume that this change can be attributed to the practice that kindergartens accept children of this age group, suggesting that local childcare availability plays a crucial mediating role in mothers' return to the labour market.

In addition to the end-December cutoff date, we also test the end-February date empirically. In our primary specification we use a 3-month time window around the December cut-off point, and define the instrument T as the following:

$$T_{iyd} = \begin{cases} 0 \; if \; 30 \leq a_{iy} \leq 32 \;\; or \; equivalently, if \;\; January \leq b_{iy} \leq March \\ 1 \; if \; 33 \leq a_{iy} \leq 35 \;\; or \; equivalently, if \;\; October \leq b_{iy} \leq December \end{cases}$$

Where a_i shows the age of the child on 1st October, expressed in months and b_i captures the month of birth of the smallest child.

The 3- month window around the end-February cutoff is as follows:

$$T_{iyf} = \begin{cases} 0 \; if \; 28 \leq a_{iy} \leq 30 \;\; or \; equivalently, if \;\; March \leq b_{iy} \leq May \\ 1 \; if \; 31 \leq a_{iy} \leq 33 \;\; or \; equivalently, if \; December \leq b_{iy} \leq Febr \end{cases}$$

Note that mothers are observed before their children turn three, hence the termination of the 3-year-long maternal benefit does not influence their labour supply decisions.

Following Lovász and Szabó-Morvai (2019), we estimate the following equations. Equation 1 refers to the reduced form estimation, where the labour market outcome of mothers is regressed directly against T and covariates.

$$Y_{iy} = \beta_y T_{iy} + X'_{iy}\Pi + \gamma_{ry} + \epsilon_{iy}$$

X_{iy} captures observables of the mother and the father[14], and γ_{ry} refers to micro-regional (LAU2) fixed effects to capture differences in local labour markets. Besides examining the effect on employment, we also consider nursery or kindergarten attendance as an outcome (Y_{iy}). This latter one gives us compliance information, meaning how many children attend kindergarten if those places are unofficially available.

In the 2-stage least square instrumental variables specification, we estimate the local average treatment effect of childcare availability on maternal employment by instrumenting the former with the child's age being less or more than the informal cut-off age. In the first stage, we regress the coverage rate in the catchment area of the mothers (see Section 4.) against the T instrument:

[14] Control variables for the mother included the following: age group, education, number of children, municipality type, and whether the mother lives in the municipality with the nursery. For the father we included age group, education, labour market status, and whether the father is present in the family.

$$C_{iyc} = \beta_y T_{iy} + X'_{iy}\Pi + \gamma_{ry} + \epsilon_{iy}$$

Where the coverage rate in the catchment area of the childcare is constructed in such a way that it assigns the kindergarten coverage rate to mothers with the smallest child being above the eligibility threshold (T_{iy}=1), while the much lower nursery coverage ratio is designated to mothers with children below the threshold:

$$C_{iyc} \stackrel{\text{def}}{=} p_{nyc}(1 - T_{iy}) + p_{kyc}T_{iy}$$

Where p_{nyc} and p_{nyc} refer to nursery and kindergarten coverage ratios in the given catchment area respectively.

The rationale behind the above definition is that while younger children can only access the local capacity of nurseries, for older children kindergartens are also available, mechanically imposing a jump from the control to the treated group in the childcare availability, resulting in a strong first stage by construction.

The second stage regresses the mother's employment status on the predicted values for the first stage:

$$Y_{iy} = \theta_y \hat{C}_{iyc} + X'_{iy}\Pi + \gamma_{ry} + \epsilon_{iy}$$

Our exclusion restriction relies on the assumption that the jump in childcare availability at the age cut-off is the only channel through which otherwise similar mothers' employment is affected by the child's age around the cut-off (having a slightly older or younger child is as good as random with regards to potential employment rate).

Using the fact that we can also directly observe childcare attendance we can also decompose the effect of childcare availability into parts. First, we can estimate the effect of childcare availability on *actual* childcare attendance following the same identification strategy as above, and regress formal childcare attendance in the second stage of the instrument.

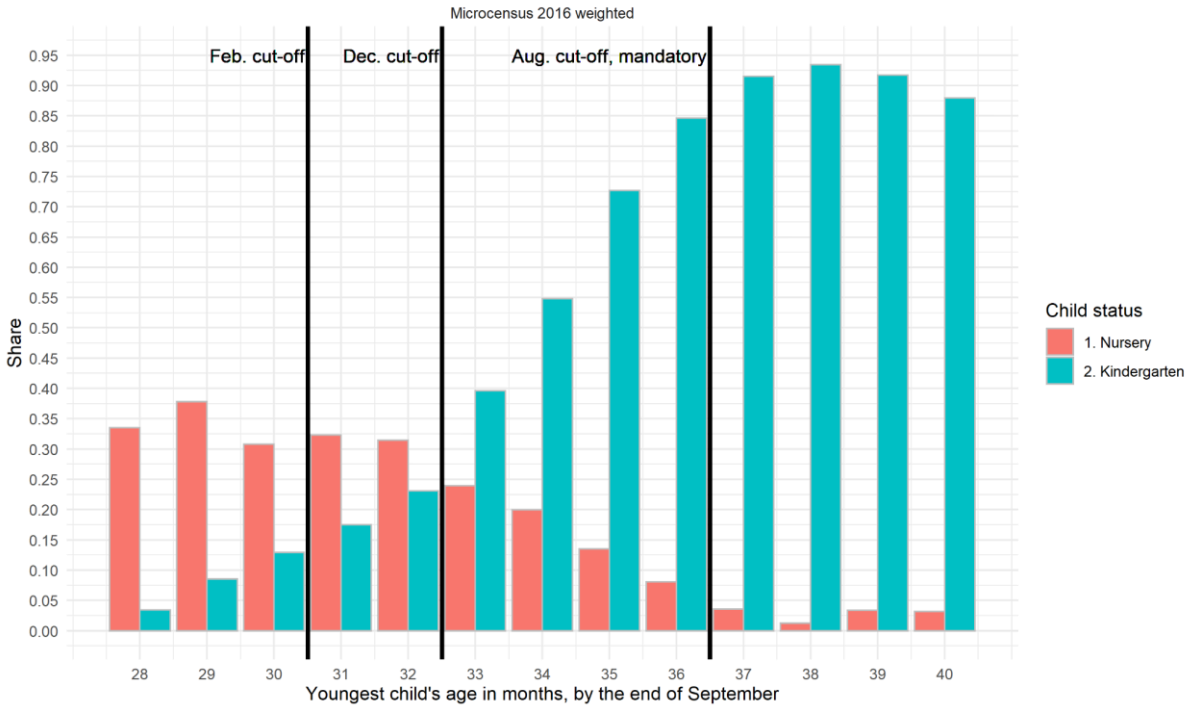
Second, it is also possible to estimate the effect of the child attending formal childcare on the mother's employment, when we use the same instrumental variable strategy, thus we instrument childcare attendance with the date of birth of the child. Hence, in the first stage, we regress the dummy variable indicating whether the child is attending (formal) childcare on the treatment indicators (as defined by the child's month of birth). Then, in the second stage, we use the predicted childcare attendance and regress it on the mother's employment status.

$$A_{iy} = \theta_y \hat{C}_{iyc} + X'_{iy}\Pi + \gamma_{ry} + \epsilon_{iy}$$

And the second step:

$$L_{iy} = \theta_y \hat{A}_{iyc} + X'_{iy}\Pi + \gamma_{ry} + \epsilon_{iy}$$

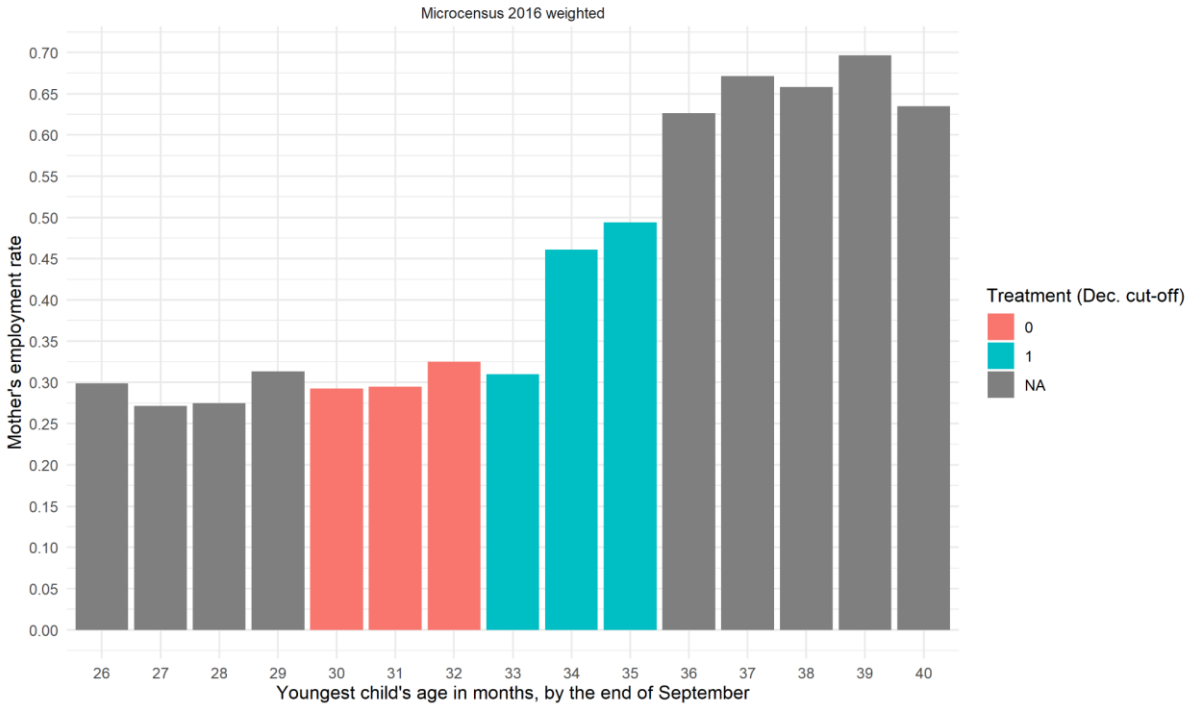
Figure 4: Youngest children's childcare facility attendance by child's age in months



Notes: The authors' calculations based on the 2016 Microcensus of Hungary.

The comparison of the treatment and control groups reveals no significant differences between the treatment and control group, except that treated mothers are slightly older (by 0,6-0,7 years) (See Table A.7. in the Appendix)

Figure 5: Maternal employment rate by youngest children's age in months according to treatment



Notes: The authors' calculations based on the 2016 Microcensus of Hungary.

7. Results

7.1. Main results

We show the results for the IV regressions in Table 3, adding more control variables across the different specifications, and the results for the reduced-form regressions can be found in Appendix A5).[15] We find that in our main specification an around 100 percentage point increase in childcare availability leads to around 7.3 percentage points higher maternal employment in 2016.[16] It is worth noting that this level of discrepancy in childcare availability is not unrealistic, as we saw earlier that on average, the jump between nursery and kindergarten coverage rates was around 120 percentage points, from 17% to about 139%, according to the definitions introduced earlier. We can interpret the estimate the following way: if mothers who can only take their young children to nurseries would have similar coverage to the level of kindergartens, their employment would increase from around 29 per cent to around 36 per cent. This is around 25% higher than the baseline employment rate, and it is around 25% of the difference between the employment rate of a mother with a child of mandatory kindergarten attendance age and one with a child 2-2.5 years of age.

[15] We also present descriptive statistics on the most crucial background variables. We only find significant differences across the treatment and the control group in terms of mothers' age, mothers in the treatment group are slightly older, by 0.5 years).

[16] We can see that adding control variables do not alter our results significantly.

Table 3: Regression results: Effect of childcare availability on maternal employment, 2016

	Outcome: Maternal employment			
	(1)	(2)	(3)	(4)
Childcare availability	0.0852***	0.0834***	0.0781***	0.0726***
	(0.0191)	(0.0176)	(0.0170)	(0.0177)
Constant	0.292***	0.161***	0.129*	0.0378
	(0.0189)	(0.0595)	(0.0755)	(0.0763)
Observations	2,505	2,505	2,505	2,505
R-squared	0.012	0.104	0.130	0.204
mean of outcome	0.3589			
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Notes: Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

This still lags behind the employment rate of around 60% of mothers whose children reached the legally mandated age of 3 years old before September. One potential reason is that even if there is the availability of childcare, not all mothers want to or can return at this point to the labour market. It can be the case due to preferences regarding formal childcare, the availability of flat maternity leave until the age of 3, already expecting another child, or facing (or perceiving) an unwelcoming labour demand amongst others. The Microcensus gives us a unique opportunity to study the channel of transmission from the availability of childcare to the reemployment of mothers. In the first step, we estimate the effect of childcare availability on children's childcare facility attendance: the regression results of this exercise are displayed in Table 4. We find that a 100-percentage point increase in childcare coverage would result in around a 17.7 percentage point increase on average in childcare attendance (with a standard error of 2.0). To put the magnitude into context, we can relate the estimate to the approximately 55% share of children not yet in childcare at the age of around 2.5 years old, that is, the growth is nearly 40%. With full childcare coverage, around one in three of the children remaining still at home could be placed in a nursery or kindergarten. For the end-February cut-off, the estimated impact is much smaller and insignificant (see Table A.6. in the Appendix).

Table 4: Regression results: Effect of childcare availability on attendance

	Outcome: Childcare facility attendance			
	(1)	(2)	(3)	(4)
Childcare availability	0.192***	0.189***	0.186***	0.177***
	(0.0210)	(0.0201)	(0.0201)	(0.0202)
Constant	0.475***	0.332***	0.269***	0.361***
	(0.0198)	(0.0554)	(0.0805)	(0.0779)
Observations	2,505	2,505	2,505	2,505
R-squared	0.059	0.116	0.125	0.202
mean of outcome	0.6263			
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Notes: Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

The second step of this exercise is to ask the question: if children do attend formal childcare before age 3 (thanks to an increase in coverage), what proportion of their mothers can find a job? Thus, we can reframe the problem from examining the effect of childcare availability to the effect of actual childcare attendance on maternal employment. Using the same assumption as before that the child born just before or after the December 31 cut-off date is as good as random with regards to potential maternal employment, we can ask how much the probability of maternal employment changes in response to being able to use childcare facilities. Table 5 presents the estimates. Being able to use childcare facility attendance would increase maternal employment probability by around 41 percentage points, conditional on maternal, and paternal observable characteristics, and the micro-region (LAU2) level fixed effects capturing local labour market circumstances. If we use the 60% employment rate of mothers with a child of kindergarten mandatory age, we could interpret the result such that being able to place the child in childcare facilities accounts for around two-thirds of the employment of mothers.

The combined estimated impact of capacity on attendance and the impact of attendance on employment are consistent with our previous estimates. Higher childcare capacity would elevate the level of attendance by around 18 percentage points, and around 40 percentage points higher share of these mothers would be able to find more employment, which would then indeed lead to an around 7 percentage point higher employment rate due to higher capacity.

Table 5: Regression results: Effect of childcare facility attendance on maternal employment

	Outcome: Maternal employment			
	(1)	(2)	(3)	(4)
Childcare attendance	0.442***	0.442***	0.421***	0.411***
	(0.0964)	(0.0890)	(0.0891)	(0.0965)
Constant	0.0818	0.0140	0.0162	-0.110
	(0.0619)	(0.0684)	(0.0821)	(0.0891)
Observations	2,505	2,505	2,505	2,505
R-squared	0.176	0.231	0.256	0.323
mean of outcome	0.3589			
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

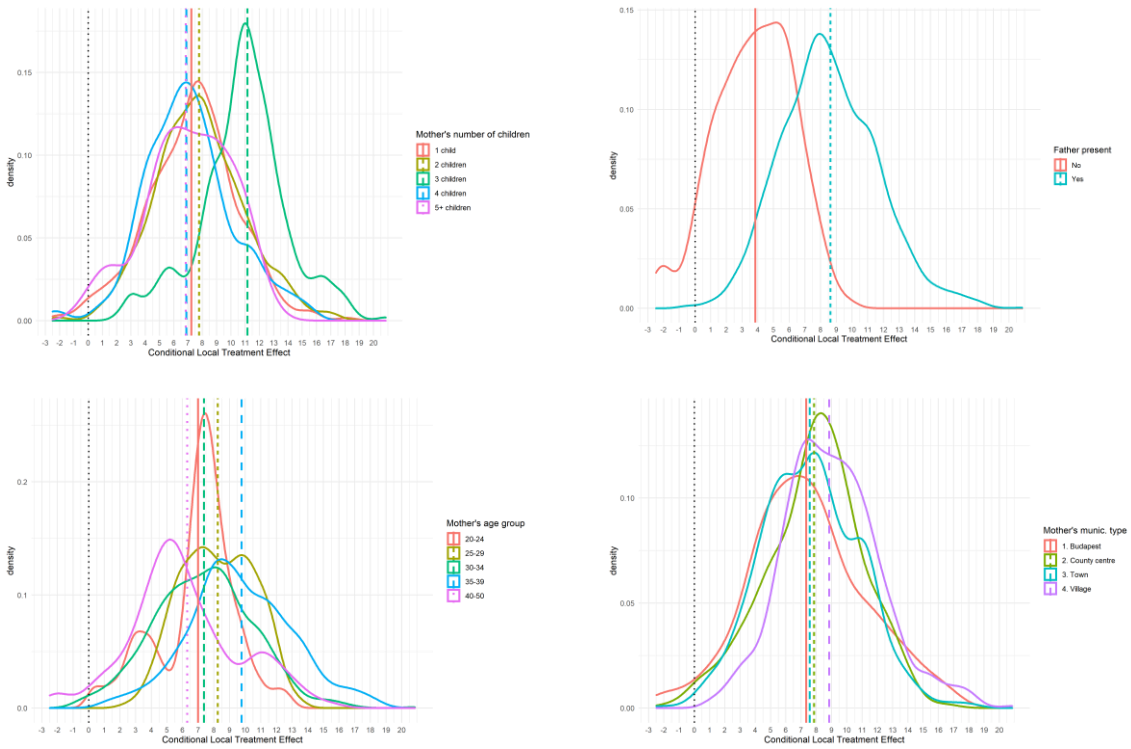
7.2. Effect heterogeneity

We re-estimate the local average treatment effect of childcare availability on maternal employment using a causal forest for instrumental variables (Athey et al., 2019). The approach uses random forests developed in the machine learning literature by Breiman (2001), to identify treatment effects in a potential outcome framework. The main advantage of estimating a causal forest is that relevant dimensions of effect heterogeneity are not handpicked by the researchers and are not restricted by pre-registration protocols, but they are revealed by joint non-parametric estimation of the conditional average treatment effects, with good asymptotic properties. Reassuringly, using this approach the overall point estimate for the local average treatment effect is 8.0 with a standard error of 2.0 for employment, and around 18.5 with a standard error of 2.2 for childcare attendance, which is close to the estimates based on the linear instrumental variable specification. We highlight those dimensions of heterogeneity which represent a larger difference compared to the baseline results. The results are presented using the out-of-bag predictions of the conditional treatment effects at the individual level, showing the densities of the estimated treatment effects grouped by the relevant control variables.

Figure 6 displays the kernel densities of the estimated treatment effects of childcare availability on maternal employment. The estimated densities tell us about large individual-level heterogeneity in effect sizes, sometimes even spanning to negative regions. We find that regarding maternal employment, the treatment effect seems to be higher for mothers with three children (average of 11.2), age group of 35-39 (9.2), or living in villages (8.8), while remarkably lower for those where the father is not present (3.8). The first results potentially point to an important unobserved factor affecting maternal labour market behaviour, namely, whether the woman has reached completed fertility. Those mothers who are at the end of their fertility cycle have a higher incentive to use the available childcare facilities, which might lead us to find higher treatment effects for mothers who are more likely to not want more children. Therefore, the main results regarding employment might underestimate the effect of childcare availability on maternal employment, if women wanted to return to work after they have finished their childbearing career. [17]

[17] To put it differently: when we estimate the effect for all women, we potentially include women who are already pregnant (and hence have no intention to work in the near future.

Figure 6: Estimated kernel densities of the conditional local treatment effect of childcare availability on maternal employment, based on causal forest estimates



Notes: based on the authors' causal forest estimates on the sample of youngest children in the 2016 Microcensus of Hungary.

Based on the information gathered from the non-parametric estimates we can turn to more traditional instrumental variable regressions to have an idea about the effect of childcare availability on maternal employment for different subsets of the population, the results are displayed in Table 6. As anticipated, in this setting as well we see larger effect sizes for the third child (15.6) and smaller municipalities without nurseries (9.4), while there is no effect for mothers to whom we cannot find the father in the family. We also cannot see large differences in whether the direct surroundings, the census counting district of the child would be considered high or low employment in the 2011 Census[18].

[18] Based on the first and third tertile of Census counting districts with regards to employment rate of 15-59 year old population. A census counting district contains 150-300 individuals.

Table 6: Heterogeneity with regressions: Effect of childcare availability on maternal employment

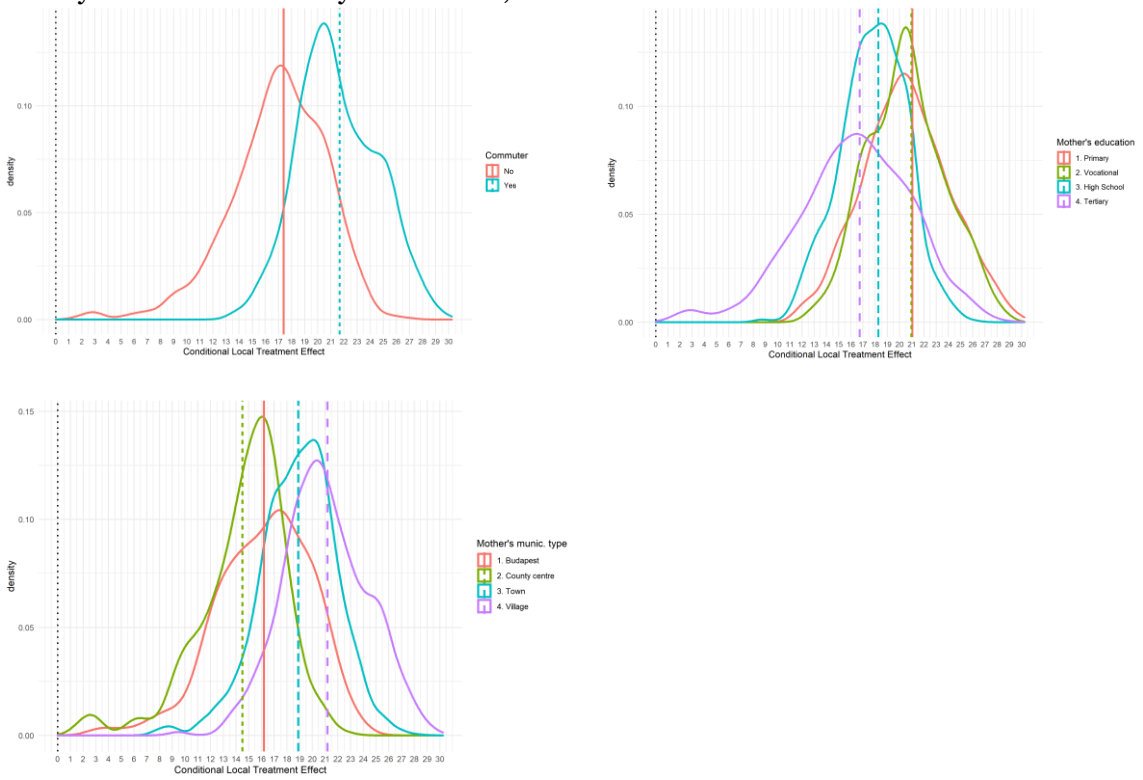
	Outcome: Maternal employment						
Subsample	Munic. with nursery	Munic. without nursery	3rd children	Towns and villages	No father present	Empl. high in 2011	Empl. low in 2011
Childcare availability	0.0514**	0.0942***	0.156***	0.0853***	-0.00708	0.0724*	0.0791***
	(0.0237)	(0.0282)	(0.0535)	(0.0182)	(0.0598)	(0.0391)	(0.0245)
Constant	0.0192	-0.140	0.504**	-0.301***	0.0970	1.046**	0.0321
	(0.0997)	(0.130)	(0.225)	(0.0918)	(0.197)	(0.433)	(0.432)
Observations	1,616	889	417	1,896	300	755	912
R-squared	0.221	0.365	0.592	0.209	0.593	0.304	0.396
mean of outcome	0.3851	0.2883	0.2444	0.3181	0.4036	0.2658	0.3982
mean of childcare availability	0.8055	0.7416	0.7195	0.7627	0.8013	0.7330	0.8077
controls	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE, counting district	mother, father, micro-region (LAU2) FE, counting district

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary with the following subsampling: municipalities with nurseries, without nurseries, only third-born children, only inhabitants of small towns and villages, no father present in the family, census counting districts in the highest tertile w.r.t. employment, census counting districts in the lowest tertile w.r.t. employment. Counting district level controls for the latter include the share of employed, ethnicity, and education levels from the previous, 2011 Census.

*** p<0.01, ** p<0.05, * p<0.1

With regards to childcare facility attendance, the densities of the conditional treatment effects are displayed in Figure 7. We find a higher effect size of around 21.6 on average for mothers who do not live in the centre of their nursery catchment area (named 'commuter' in the figure), have primary or lower secondary educational attainment, live in a village, or are less than 29 years old (each with an around 21.0 local average treatment effect). We can see that there is substantial heterogeneity in terms of the effect of additional childcare facility places on realised attendance, this channel being more important for mothers with a disadvantaged socioeconomic background. The higher effect for mothers who live in municipalities without a nursery suggests that it might be the case that this effect heterogeneity originates from the changing proximity (childcare being locally available) rather than the increased number of places in the area.

Figure 7. Estimated kernel densities of the conditional local treatment effect of childcare availability on childcare facility attendance, based on causal forest estimates



Notes: based on the authors' causal forest estimates on the sample of youngest children in the 2016 Microcensus of Hungary.

We can estimate regression equations for childcare facility attendance as well. Again, we see more pronounced effects of childcare availability for smaller municipalities without nurseries (0.28), and at the third child (0.31), but there are somewhat different magnitudes in census counting districts[19] with high employment vs. the ones with low employment (see Table 7). For children where employment is lower, higher childcare availability would induce around 23 percentage points higher attendance, around 28% higher than the baseline of around 18 pp., while we condition for the micro-regional fixed effects and the municipality type as well. This suggests again that the availability of childcare facilities has more impact in areas of the country with fewer opportunities in terms of facility attendance, even if it does not necessarily turn into more employment at the date of observation.

[19] These districts typically contain around 1000 persons, hence the employment measure used here is not a characteristic of the local labour market, rather, it is a social composition index.

Table 7: Heterogeneity with regressions: Effect of childcare availability on attendance

	Outcome: Childcare facility attendance						
subsample	Munic. with nursery	Munic. without nursery	3rd children	Towns and villages	No father present	Empl. high in 2011	Empl. low in 2011
Childcare availability	0.150***	0.280***	0.306***	0.201***	0.161***	0.153***	0.232***
	(0.0240)	(0.0334)	(0.0584)	(0.0237)	(0.0575)	(0.0371)	(0.0282)
Constant	0.303**	0.316**	0.805*	0.443***	0.408**	0.860	0.713
	(0.138)	(0.124)	(0.445)	(0.104)	(0.201)	(0.587)	(0.555)
Observations	1,616	889	417	1,896	300	755	912
R-squared	0.227	0.403	0.581	0.240	0.559	0.270	0.396
mean of outcome	0.6487	0.5658	0.5502	0.6121	0.6234	0.5703	0.6441
mean of childcare availability	0.8055	0.7416	0.7195	0.7627	0.8013	0.7330	0.8077
controls	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE	mother, father, micro-region (LAU2) FE, counting district	mother, father, micro-region (LAU2) FE, counting district

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary with the following subsampling: municipalities with nurseries, without nurseries, only third-born children, only inh abitants of small towns and villages, no father present in the family, census counting districts in the highest tertile w.r.t. employment, census counting districts in the lowest tertile w.r.t. employment. Counting district level controls for the latter include the share of employed, ethnicity, and education levels from the previous, 2011 Census.

*** p<0.01, ** p<0.05, * p<0.1

7.3. Robustness

We performed several robustness checks to test the validity of the results. We found that the treatment effects are robust to the definition of maternal labour force status, as the estimates are quite similar whether we look at participation rate (including unemployed), employment excluding public work, or including mothers who are also students, pensioners, or are on childcare benefits; the point estimates range from 6.2 to 7.3 (see Table 8). We also estimated the treatment effects including all children, not only the youngest ones in the family. These results would suggest a lower, around 4.7 local average treatment effect with a standard error of around 1.4. This further suggests that mothers at the end of their fertility cycle react more to childcare availability. The estimates are also similar if we do not use the sample weights of the 2016 Microcensus.

Table 8: Regression results: Effect of childcare availability on maternal employment using different employment definitions and sample restrictions

Outcome:	Employed (broad) + Unemployed	Employed (broad)	Employed (without public work)	Employed (unweighted)	Employed (not only youngest children)
Childcare availability	0.0675***	0.0624***	0.0724***	0.0692***	0.0470***
	(0.0174)	(0.0185)	(0.0175)	(0.0137)	(0.0142)
Constant	0.259***	0.264***	0.0625	0.0321	0.0444
	(0.0746)	(0.0762)	(0.0735)	(0.0577)	(0.0611)
Observations	2,505	2,505	2,505	2,505	3,136
R-squared	0.207	0.220	0.216	0.210	0.184
controls	mother, father, micro-region (LAU2) level units FE				

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. Regressions with different definitions of employment, weights, and sample, the first four columns are based on the sample of youngest children within families.

*** p<0.01, ** p<0.05, * p<0.1

We experimented with 2-month and 4-month time windows around the cut-off as well and checked the alternative cut-off at the end of February. The 4-month cut-off would result in a higher, around 12 percentage point effect, but there can be a reason for concern that treatment and control groups are not as comparable as some children end up with 8 months of age difference in the sample (Table 9). Regarding the 2-month window which would be the strongest with regards to identification, depending on the set of controls regressions yield around a 3.5-5 percentage point effect, however potentially due to the low number of observations the estimates are not always statistically significant at the 5% level.

Table 9: Regression results: Effect of childcare availability on maternal employment using different cut-offs and time windows

	Outcome: Maternal employment					
Cut-offs	End of December			End of February		
Time window	2 months	3 months	4 months	2 months	3 months	4 months
Childcare availability	0.0327	0.0726***	0.123***	0.0105	0.0222	0.0503***
	(0.0236)	(0.0177)	(0.0156)	(0.0187)	(0.0148)	(0.0138)
Constant	-0.0107	0.0378	-0.0931	0.103	0.0543	-0.0692
	(0.100)	(0.0763)	(0.0595)	(0.0913)	(0.0717)	(0.0578)
Observations	1,696	2,505	3,449	1,690	2,566	3,479
R-squared	0.238	0.204	0.182	0.219	0.186	0.171
mean of outcome	0.34	0.36	0.39	0.31	0.30	0.31
Controls	mother, father, micro-region (LAU2) FE					

Notes: Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using different time windows and cut-offs for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

Finally, we implement placebo exercises to validate the results, displayed in Table 10. We shifted the cut-off dates plus-minus 6 months in terms of children's age to test whether our estimates are just an artefact of a more general trend of returning to the labour market, or indeed there is a special significance of the informal December cut-off. Reassuringly, we find no statistically significant estimates at the 5% level along the pseudo-cut-offs, supporting our assumption that the age around the cut-off value could be considered as good as random concerning potential employment.

Table 10: Regression results: Placebo effect of childcare availability on maternal employment

	Outcome: Maternal employment			
Placebo cut-off	End of June (younger children)		End of June (older children)	
Childcare availability	0.0272	0.0284*	0.00318	0.0203
	(0.0169)	(0.0170)	(0.0196)	(0.0171)
Constant	0.235***	0.171***	0.648***	-0.105
	(0.0150)	(0.0556)	(0.0192)	(0.0812)
Observations	2,956	2,956	2,388	2,388
R-squared	0.002	0.158	0.000	0.270
mean of outcome	0.25	0.25	0.65	0.65
Controls	none	mother, father, micro-region (LAU2) FE	none	mother, father, micro-region (LAU2) FE

Notes: Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the placebo cut-offs (±6-month shifting away from the informal December 31 cut-off for birthdate) as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary.
*** p<0.01, ** p<0.05, * p<0.1

7.4. Comparison with 2011

We executed the same analysis as before on data from the 2011 Census, which we briefly present here. [20] We show three specifications (with all control variables included). Our main finding is that a 100 percentage points increase in coverage would lead to an increase in the employment rate by about 4.7 percentage points, which is only two-thirds of the estimated parameter in 2011 (see Table 11).

Table 11: Regression results: Effect of childcare availability on maternal employment, 2011

	Outcome: Maternal employment			
Childcare availability	0.0449***	0.0465***	0.0470***	0.0468***
	(0.00403)	(0.00393)	(0.00398)	(0.00402)
Constant	0.299***	0.0653***	0.0977***	-0.0320*
	(0.00804)	(0.0183)	(0.0197)	(0.0169)
Observations	34,808	34,808	34,808	34,808
R-squared	0.003	0.133	0.146	0.156
mean of outcome			0.3305	
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2011 Census of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

This lower estimate relative to 2016 does not seem to be related to the lower estimated impact of coverage ratio on childcare attendance, which is estimated to be slightly higher, with 19.2

[20] It is worth mentioning that since sample sizes are 10 times larger (we observe the universe of mothers), all regression estimates tend to be statistically significant.

percentage points more children attending formal childcare, when it becomes available (see Table 12).

Table 12: Regression results: Effect of childcare availability on Childcare facility attendance, 2011

	Outcome: Childcare facility attendance			
Childcare availability	0.173***	0.175***	0.174***	0.175***
	(0.00617)	(0.00565)	(0.00562)	(0.00569)
Constant	0.420***	0.297***	0.287***	0.273***
	(0.0102)	(0.0216)	(0.0277)	(0.0232)
Observations	34,808	34,808	34,808	34,808
R-squared	0.043	0.119	0.122	0.140
mean of outcome			0.5415	
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2011 Census of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

However, the effect of the child attending (formal) childcare on mothers' employment is substantially lower, with the estimated effect being only 27 percentage points, which is only two-thirds of the estimated coefficient from 2016 (see Table 13). One possible explanation lies in the different macroeconomic circumstances: 2011 marked the lowest point of employment throughout the 15 years examined here, thus, women with young children returning to the labour market had significantly lower chances of taking up employment than five years later (when the labour market prospects of women were rapidly improving).

Table 13: Regression results: Impact of childcare facility attendance on mothers' employment

	Outcome: Maternal employment			
Childcare facility attendance	0.259***	0.266***	0.270***	0.268***
	(0.0220)	(0.0223)	(0.0227)	(0.0228)
Constant	0.190***	-0.0138	0.0205	-0.105***
	(0.0126)	(0.0181)	(0.0202)	(0.0184)
Observations	34,808	34,808	34,808	34,808
R-squared	0.156	0.243	0.255	0.262
mean of outcome			0.3305	
sd of outcome			0.4704	
share of treated			0.4947	
sd of share of treated			0.5000	
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2011 Census of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

7.5 Comparison with results based on the Labour Force Survey of Hungary

We implement the same empirical strategy on a different data source, the Labour Force Survey of Hungary, using a pooled sample of 4th quarters in the period 2005-2020, with the same controls and instrument used for the Microcensus samples. We use the ILO definition of employment as the outcome variable: all those of working age who either worked for at least one hour during the reference week or had a job but were absent due to sick leave, maternity leave, etc. By analysing LFS we can observe both the period of recession and the period of expansion. We investigate mothers at different points in time, in September, October, November and December with the same child age. When defining control and treatment groups, the range of the youngest child's age for the control group includes 30–35-month-olds and the treatment group includes 33–38-month-olds. Children who turn 3 before the cut-off point make up the treatment group, while those turning 3 after the cut-off form the control group (see Table 14). The children of mothers in our treatment group (those born before January 1st) may enrol between September and January, while control group mothers are forced to wait until the next school year. The end-of-February cut-off focuses on those children who reached the age of 2.5 and may have been admitted to kindergarten in September of the previous year depending on available excess capacity. In this case, the treatment group represents those born prior to the end of February, while the control group covers children who do not reach the age of 3 by the cut-off date (see Table 14).

Table 14: Defining treatment and control groups for December and February cut-offs using a 3-month window

	Age of the youngest child (in months)			
Cut-off	End of December		End of February	
Month of observation	Treatment group	Control group	Treatment group	Control group
9	35, 34, 33	32, 31, 30	33, 32, 31	30, 29, 28
10	36, 35, 34	33, 32, 31	34, 33, 32	31, 30, 29
11	37, 36, 35	34, 33, 32	35, 34, 33	32, 31, 30
12	38, 37, 36	35, 34, 33	36, 35, 34	33, 32, 31

Notes: We consider 2 cut-off dates, the end of December and the end of February with a 3-month time window. Additionally, we include a dummy variable in our regressions to control for children aged over 3.

In our primary specification, we control for observables of the mother and the father, along with microregion regional fixed effects. Control variables for the mother included the following: age group, education, number of children and municipality type (capital city and county capitals, cities, towns). For the father, we included age, education and the labour market status. As shown in Table 15-16, we find that using this alternative data source yields similar results to that of the Microcensus.

Main results

We show the results for the reduced form and IV regressions in the tables below (see Tables 15-16.). We find that in our main specification (with the end of December cut-off using a 3-month time window) a 100 percentage points increase in childcare availability leads to around 6 percentage points higher maternal employment rate in the period between 2008 and 2020. While the corresponding reduced form equation indicates that the maternal employment rate is 7 percentage points higher if mothers are eligible for kindergarten rather than a nursery school. There is a significant positive effect of childcare availability on the maternal employment rate during the expansion period (2016-2019) showing a 10.5 percentage point estimated effect. This is largely in line with previous results from the Microcensus and the Census, hence we can corroborate that the availability of childcare has a more positive effect on female employment

when the demand for labour is higher. The estimation using the February cut-off yields smaller and less significant results than we expected (see Appendix).

Table 15: Reduced form regression results with the end-of-December cut-off, LFS, 2005-2020

	Outcome: Maternal employment				
Specification	Reduced form				
Cut-off	End of December				
	(1)	(2)	(3)	(4)	(5)
Time period	2008-2020	2005-2009	2010-2015	2015-2019	2016-2019
Treatment	0.0693***	-0.00272	0.0538	0.119***	0.126**
	(0.0250)	(0.0345)	(0.0358)	(0.0447)	(0.0508)
Constant	0.0618	0.0516	-0.0199	0.148*	0.193*
	(0.0585)	(0.0696)	(0.078)	(0.0883)	(0.0987)
Observations	3,939	1,687	971	930	1,115
R-squared	0.230	0.275	0.342	0.393	0.421

Notes: Robust standard errors in parentheses clustered at the járás (micro-region) level.

*** p<0.01, ** p<0.05, * p<0.1

Table 16: 2SLS regression results with the end of December cut-off, 2005-2020

	Outcome: Maternal employment				
Specification	2SLS estimation				
Cut-off	End of December				
	(1)	(2)	(3)	(4)	(5)
Time period	2008-2020	2005-2009	2010-2015	2015-2019	2016-2019
Childcare availability	0.0592***	-0.00242	0.0468	0.0986***	0.105***
	(0.0207)	(0.0287)	(0.0294)	(0.0342)	(0.0382)
Constant	0.332	-0.138	-0.306**	0.540***	0.606***
	(0.216)	(0.104)	(0.123)	(0.147)	(0.142)
Observations	3,939	1,687	971	930	1,115
R-squared	0.228	0.275	0.342	0.392	0.418

Notes: Robust standard errors in parentheses clustered at the járás (micro-region) level.

*** p<0.01, ** p<0.05, * p<0.1

8. Discussion and Conclusions

We used a clean identification strategy, building on the date of birth of children and the rules governing admission to kindergarten to estimate the effect of increasing the coverage of (formal) early childhood care on mothers' labour supply. Our results show that expanding nurseries to full capacity would likely increase mothers' employment rate by about 25 per cent, for mothers with children between the ages of 2-3. This effect is similar to that found in Lovász and Szabó-Morvai (2019), but we need to emphasize that it has been estimated in a period following a substantial expansion of early childcare availability. In other words, the effect is estimated on a sample of women who are *not* the keenest to place their children in formal childcare (as they likely also found a way to place their children in childcare before the expansion). It is also worth noting that it is an intention to treat effect for the whole population of mothers with children aged between 2 and 3, and we cannot exclude from the sample those who are already pregnant, and hence have no intention of returning to the labour market. It is also worth pointing out that mothers with children under age 3 are still eligible for monetary childcare benefits, which also entitle these mothers to social security coverage. Thus, the effect is estimated in an environment where mothers do not have strong monetary incentives to return to work and in a social environment where it is largely accepted to be out of work for 3 years after childbearing.

Our data also permitted us to gauge to what extent the potential availability of childcare leads to actual nursery attendance. Our estimates suggest that full childcare availability would lead to the attendance of around one-third of children not inscribed in formal childcare at around age 2.5. We find substantial heterogeneity in the usage of potential childcare: those living in municipalities with no nurseries have much higher effects, as well as those who have three children and presumably have completed their childbearing.

Our findings are informative of heterogeneities in the effect of childcare availability on maternal employment. Yet again, among those with 3 children, and among those living in villages or municipalities without nurseries (these two largely overlap), the effects are larger. This squares well with the effects of availability on childcare attendance, hence we conjecture that the direct effect of having a child attending childcare on employment is similar in all groups. This latter effect is very pronounced and amounts to 41 percentage points. A comparison with 2011, when the labour market was slack, reveals that the effect of the child attending formal childcare on maternal employment is dependent on the economic cycle: this was much more muted in 2011.

The findings above point toward a topic of research which is also very relevant from a policy perspective. Given the relatively low effect of childcare availability on attendance, it would be useful to conduct further research into this issue. Most notably, one would what to know whether this relatively low take-up is due to (i) a lack of information on the possibility of taking children below age 3 to kindergarten; (ii) the actual (or perceived) attitude of kindergartens towards admitting young children; (iii) the availability of alternative, informal forms of childcare.

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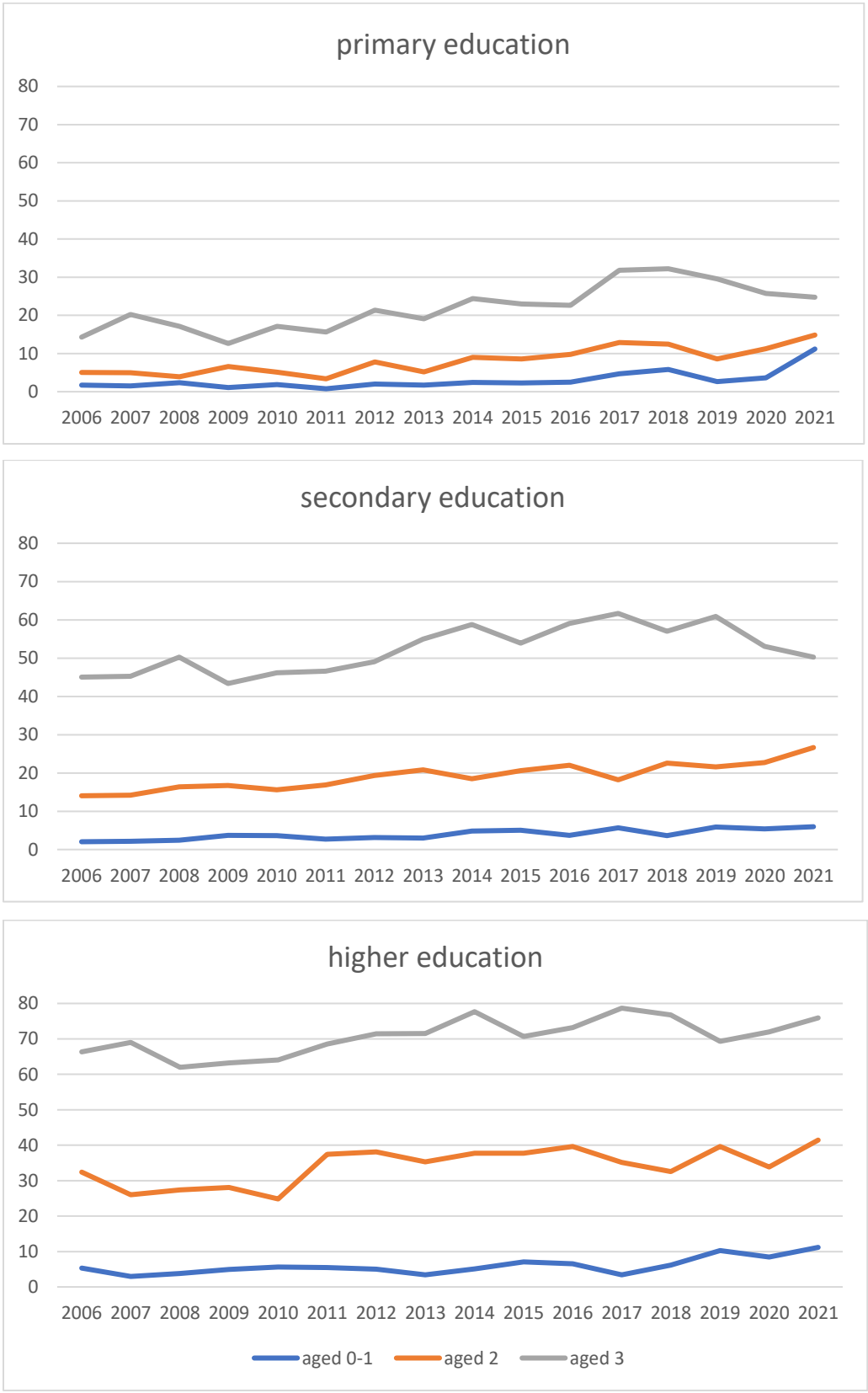
Appendices

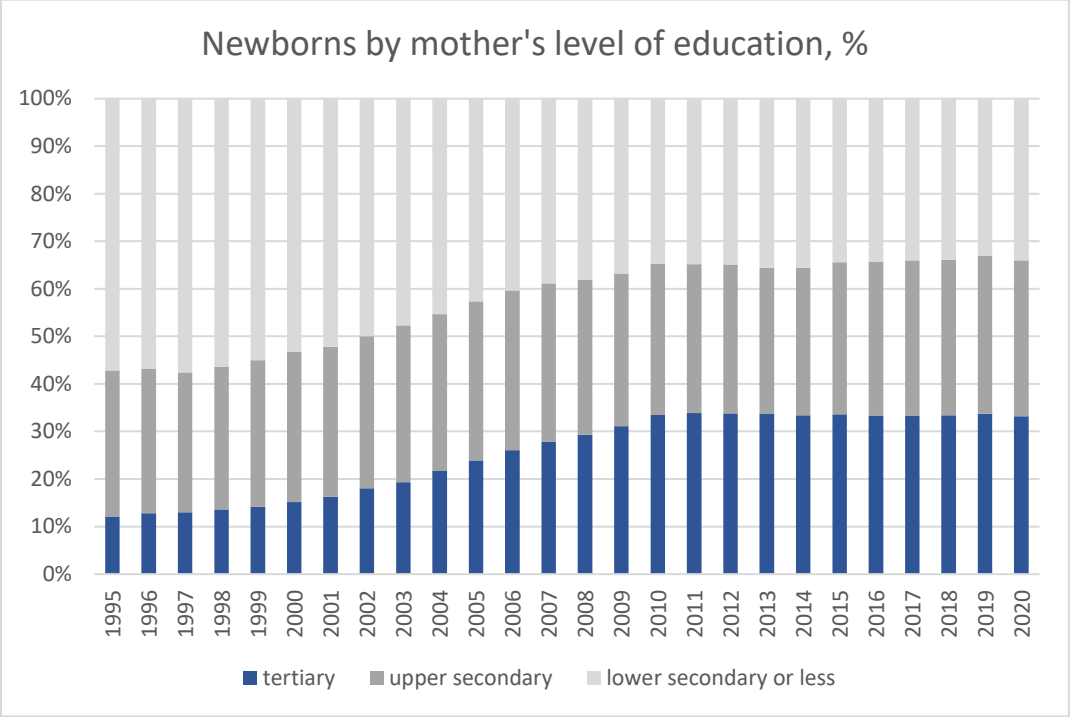
A.1 Background information on maternal employment and education levels

Figure A.1: Time trend in maternal employment by level of education and age of the youngest child



The figures above are based on LFS weighted data (employed defined as worked 1 hour).

Figure A.2: Time trend in mothers' employment level



Source: Administrative data reported to the Central Statistical Office (KSH) by hospitals.

A.2. Rules affecting access to childcare[21]

By default, access to nurseries is conditional on the parents' employment, but municipalities can deviate from the centrally provided guidelines in this respect.

Family day-care centres[22]

Family day-care centres were introduced as a form of daycare in 1997 and are eligible for a per capita public subsidy since 2003. They gained significant ground between 2005-2016, providing additional daycare capacity to compensate for the shortage of nursery places in both large and small towns. Family day care centres need to obtain a licence, but requirements are easier to fulfil both regarding the infrastructure and staff qualifications. Service providers are required to have at least 8 years of primary education and complete a 40-hour course (in practice most providers have higher qualifications and prior experience in public childcare). The centres were intended for children from the age of 20 weeks to 14 years, however, the age limit was reduced to 3 years in 2017.

Kindergarten admission before age of three

Between September 2009 and August 2017, kindergartens were allowed to enrol nursery-age children in a mixed-age group. The so-called unified (or integrated) nursery-kindergartens could operate in villages and towns of below 10 000 inhabitants where there were free capacities in the kindergarten, and there was no nursery.[23] There could be one such group of mixed ages and admit a maximum of five children aged over 2 years. This format was phased out in August 2017.

[21] ECEC is mainly regulated by the 1997 Act in the Protection of Childcare. Other relevant legislations include the 1993 Education Act (amended in 2011), the Government Decree 229/2012 on the implementation of the Education Act, the Ministerial Decree 20/2012 on the management of educational institutions and the Ministerial Decree 15/1998, IV.30 about the Task and Operational Criteria of Child Protection Institutes and Personnel.

[22] https://www.ksh.hu/docs/hun/xftp/idoszaki/pdf/kisgyermnapkozbeni.pdf

[23] https://www.koloknet.hu/ovoda/egyseges-ovoda-bolcsode/

At the same time, a similar, new format was introduced: setting up a mini-nursery, kindergartens can start a small group (up to 8) for nursery-age children.[24]

As of September 2011, kindergartens may admit a child who will reach the age of three within six months of admission. The child must be at least 2.5 years old on the day of admission.[25]

Introduction of care fees in nurseries[26]

As of 2012, municipalities can charge a fee for daycare in nurseries to cover the difference between the actual cost and the public subsidy. The fee is affordable to most families as it must be proportional to the parents' monthly income.[27]

Compulsory kindergarten

Since 2015, childcare from the age of three is obligatory, however, the obligation can be waived in justified cases (e.g. if the child has a severe disability or illness).

Municipal obligations for providing childcare

As of 2017, the child protection law obliges local governments to regularly assess needs and provide daycare for children aged below 3 if there are 40 or more children below the age of three living in the settlement, or if at least five parents request the service. The actual enforcement of this rule was postponed first until the end of 2018, then to the end of 2020. Previously, only settlements with more than 10,000 inhabitants were obliged to organise nursery services in their communities.[28]

Regulation of children/carer rates[29]

Children/carer rates are regulated by the central government. The number of carers is set by government decrees that prescribe the number of carers by group and set the maximum group size. Until 2010, the maximum number of children in a group was somewhat ambiguously defined: the relevant decree set the maximum at ten but also referred to the guidelines issued by the research institute of the relevant ministry, which stated that, if needed, nurseries may deviate by 20% from the maximum set by the law, implying a maximum of 12 children per group. This ambiguity was eliminated by the update of the guidelines in 2011, which stated that the maximum set by the law should be kept[30] – however, by that time the legal maximum had been raised to 12.[31]

A.3 Rules affecting access to parental leave benefits

Maternity leave and cash transfers in Hungary[32]
In the first six months after the birth of the child, the mother is eligible to receive the CSED (the father is eligible only if the mother died or lost her parental rights). GYED is available to both parents until the child becomes two years old and covers 70% of the previous wage, capped at 1.4 times the minimum wage. As of 2014 new rules (labelled "GYED Extra") allow the parent receiving parental leave to take up full-time employment when the child turns one. As of 2016,

[24] https://kormanyablak.hu/hu/feladatkorok/353/SZGYH00203
[25] Hungary - Act CXC of 2011 on National Public Education.
[26] http://www.ksh.hu/docs/hun/xftp/stattukor/kisgyermnapkozbeni/kisgyermnapkozbeni12.pdf
[27] The fee must not exceed 25% of the per capita net income of the family and is waived for families eligible for the "child-support allowance" (gyermekvédelmi kedvezmény), which applied to one third of children enrolled in 2012. http://www.ksh.hu/docs/hun/xftp/stattukor/kisgyermnapkozbeni/kisgyermnapkozbeni12.pdf
[28] Central Statistical Office (2018). A kisgyermekek napközbeni ellátása. https://www.ksh.hu/docs/hun/xftp/stattukor/kisgyermnapkozbeni/kisgyermnapkozbeni16.pdf.
[29] https://net.jogtar.hu/jogszabaly?docid=99800015.nm
[30] https://macske.hu/wp-content/uploads/2020/03/macske.hu-feladatok-bolcsodei-neveles-gondozas-orszagos-alapprogramja.pdf
[31] http://www.jogiportal.hu/view/15-1998-iv-30-nm-rendelet
[32] Based on Erős et al. (2022). Young women on the labour market in Hungary and Poland [Unpublished manuscript]. Youth Employment PartnerSHIP.

this rule was further relaxed allowing the return to work six months after the birth of the child while receiving the full amount of GYED.

The flat rate GYES is universally available until the child turns three regardless of previous employment. GYES also provides health care and pension insurance coverage, but only a low benefit amount of 28,500 HUF per month (around 8% of the average wage) for one of the parents. For parents with at least three children, the GYET provides a fixed amount of 28,500 HUF until the youngest child turns eight.

A.4 Summary of parental leave transfers in Hungary in 2022[33]

Age of child	Worked before	Transfer	Eligibility	Compensation rate	Flexibility/Work
0-2 month	yes	Paid paternity leave	Only the father, paid paternity leave of **5 days** in the first 2 months after birth	100% of previous earnings	-
0-6 month	yes	CSED	Only the mother, if -she has been employed at least for 365 days within the two years before the birth Or - have completed two semesters at an accredited higher education institution within two years prior to the birth	Until 2021: 70% of previous earnings From 2021: 100% of previous earnings	Employment not allowed
0-6 month	no	GYES	See GYES		
6-24 month	yes	GYED	Either of the parents living with the child, under the same conditions as CSED	70% of previous earnings, until max 140% of the statuary min. wage	Work unlimited hours after the child becomes 6 months old
6-24 month	no	GYES	See GYES		
24-36 month	-	GYES	All parents. Also foster parents and guardians (grandparents: after the child turns 1)	Fixed amount of 28,500 HUF / month (around 8% of avg. wage)	Work unlimited hours after the child turns 6 months
3-8 years	-	GYET	Either of the parents in a family with three or more children - leave during the period between the 3rd and 8th birthday of the youngest child	Fixed amount of 28,500 HUF / month	30 hours a week, or unlimited hours if the work is done at home

[33] Based on Erős et al. (2022). Young women on the labour market in Hungary and Poland [Unpublished manuscript]. Youth Employment PartnerSHIP.

A.5 Results for the reduced form regressions, 2016 Microcensus

Table A.1: Results for the reduced form regressions, 2016 Microcensus of Hungary

	Outcome: Maternal employment			
Treatment	0.103***	0.101***	0.0945***	0.0875***
	(0.0232)	(0.0213)	(0.0208)	(0.0224)
Constant	0.306***	0.177***	0.143*	0.137*
	(0.0171)	(0.0599)	(0.0764)	(0.0757)
Observations	2,505	2,505	2,505	2,505
R-squared	0.012	0.105	0.13	0.203
mean of outcome	0.36			
sd of outcome	0.48			
share of treated	0.51			
sd of share of treated	0.50			
mean of childcare availability	0.79			
sd of childcare availability	0.62			
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Notes: Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Reduced form estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups.

*** p<0.01, ** p<0.05, * p<0.1

Table A.2: Results for the reduced form regressions, 2011 Census of Hungary

	Outcome: Maternal employment			
Treatment	0.0496***	0.0515***	0.0520***	0.0517***
	(0.00450)	(0.00435)	(0.00440)	(0.00446)
Constant	0.306***	0.0745***	0.107***	0.119***
	(0.00800)	(0.0182)	(0.0196)	(0.0164)
Observations	34,808	34,808	34,808	34,808
R-squared	0.003	0.133	0.146	0.156
mean of outcome	0.33			
sd of outcome	0.47			
share of treated	0.49			
sd of share of treated	0.50			
mean of childcare availability	0.70			
sd of childcare availability	0.57			
controls	none	mother	mother, father	mother, father, micro-region (LAU2) FE

Notes: Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Reduced form estimates using a ±3-month time window around the informal December 31 cut-off for birthdate as treated and control groups.

*** p<0.01, ** p<0.05, * p<0.1

Table A.3: Comparing treated and control groups along the main variables based on the 2016 Microcensus

Variables	ß (treatment's coefficient)	se of ß	ß (treatment's coefficient)	se of ß
Cut-off	December		February	
mother age	0.578**	(0.257)	-0.0963	(0.260)
father age	0.559*	(0.314)	-0.282	(0.305)
primary educ.	-0.00401	(0.0173)	-0.0109	(0.0165)
lower secondary	-0.00419	(0.0155)	0.0306*	(0.0156)
higher secondary	0.0244	(0.0223)	0.0265	(0.0221)
tertiary educ.	-0.0163	(0.0240)	-0.0462*	(0.0240)
father dummy	-0.00747	(0.0139)	-0.0120	(0.0138)
Budapest	0.0148	(0.0203)	-0.0128	(0.0202)
village	-0.00651	(0.0218)	0.00565	(0.0217)
city with county right	0.00298	(0.0218)	0.0128	(0.0216)
town	-0.0112	(0.0204)	-0.00563	(0.0200)
N	2505		2566	

Notes: 2016 Microcensus. The table displays coefficients of a regression of the different covariates on the treatment indicator on a sample consisting of the treatment and control groups, using the baseline specification with a 3-month window.

Table A.4: Regression results: Effect of childcare availability on maternal employment with different time windows and cut-offs, 2016

	Outcome: Maternal employment					
Cut-offs	December cut-off			February cut-off		
Time window	2 months	3 months	4 months	2 months	3 months	4 months
Childcare availability	0.0265	0.0624***	0.0884***	-0.00171	0.0181	0.0406**
	(0.0251)	(0.0185)	(0.0152)	(0.0229)	(0.0189)	(0.0159)
Constant	0.271***	0.264***	0.126**	0.250**	0.164**	0.0728
	(0.102)	(0.0762)	(0.0615)	(0.1000)	(0.0783)	(0.0646)
Observations	1,696	2,505	3,449	1,690	2,566	3,479
R-squared	0.258	0.220	0.196	0.256	0.209	0.191
mean of outcome	0,4424	0,4502	0,4764	0,4162	0,4091	0,4176
sd of outcome	0,4968	0,4976	0,4995	0,4931	0,4918	0,4932
share of treated	0,5016	0,5086	0,5031	0,4836	0,4923	0,4904
sd of share of treated	0,5001	0,5000	0,5001	0,4999	0,5000	0,5000
mean of childcare availability	0,7781	0,7882	0,7836	0,7620	0,7710	0,7676
sd of childcare availability	0,6188	0,6204	0,6212	0,6218	0,6202	0,6193
Controls	mother, father, micro-region (LAU2) FE					

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using different time windows and cut-offs for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2016 Microcensus of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

Table A.5: Regression results: Effect of childcare availability on maternal employment with different time windows and cut-offs, 2011

	Outcome: Maternal employment					
Childcare availability	0.0209***	0.0468***	0.0994***	0.0201***	0.0322***	0.0467***
	(0.00520)	(0.00402)	(0.00414)	(0.00593)	(0.00487)	(0.00415)
Constant	-0.0646***	-0.0320*	-0.00483	-0.0173	0.0252	0.0311*
	(0.0213)	(0.0169)	(0.0153)	(0.0222)	(0.0188)	(0.0164)
Observations	23,234	34,808	47,051	23,177	34,890	46,932
R-squared	0.158	0.156	0.165	0.147	0.144	0.146
mean of outcome	0,3205	0,3305	0,3542	0,3018	0,2974	0,2984
sd of outcome	0,4667	0,4704	0,4783	0,4591	0,4571	0,4575
share of treated	0,4886	0,4947	0,4997	0,4970	0,5008	0,4951
sd of share of treated	0,4999	0,5000	0,5000	0,5000	0,5000	0,5000
mean of childcare availability	0,6952	0,7022	0,7078	0,7050	0,7098	0,7030
sd of childcare availability	0,5670	0,5664	0,5663	0,5654	0,5660	0,5659
Cut-offs	End of December			End of February		
Time window	2 months	3 months	4 months	2 months	3 months	4 months
Controls	mother, father, micro-region (LAU2) FE					

Robust standard errors in parentheses clustered at the micro-region (LAU2) level. Instrumental variable estimates using different time windows and cut-offs for birthdate as treated and control groups. The sample contains the youngest children in the family unit of the 2011 Census of Hungary.

*** p<0.01, ** p<0.05, * p<0.1

A.6 Comparison with the Hungarian Labour Force Survey: regression results using the end of February cut-off

Table A6: Reduced form regression results with the end of February cut-off, 2005-2020, LFS

	Outcome: Maternal employment				
Specification	Reduced form				
Cut-off	End of February				
	(1)	(2)	(3)	(4)	(5)
Time period	2008-2020	2005-2009	2010-2015	2015-2019	2016-2019
Treatment	0.0411*	0.0245	0.0574*	0.0544	0.0605
	(0.0215)	(0.0284)	(0.0296)	(0.0369)	(0.0437)
Constant	0.0724	0.0137	-0.0642	0.215**	0.323***
	(0.0558)	(0.0677)	(0.0698)	(0.0976)	(0.112)
Observations	3,894	1,663	989	911	1,099
R-squared	0.190	0.270	0.359	0.360	0.387

Notes: Robust standard errors in parentheses clustered at the járás (micro-region) level.

*** p<0.01, ** p<0.05, * p<0.1

Table A7: 2SLS regression results with the end of February cut-off, 2005-2020, LFS

	Outcome: Maternal employment				
Specification	2SLS estimation				
Cut-off	End of February				
	(1)	(2)	(3)	(4)	(5)
Time period	2008-2020	2005-2009	2010-2015	2015-2019	2016-2019
Childcare availability	0.0357**	0.0216	0.0494**	0.0453	0.0507
	(0.0181)	(0.0233)	(0.0240)	(0.0283)	(0.0330)
Constant	-0.0213	-0.151	-0.344***	0.0935	0.176
	(0.0821)	(0.116)	(0.105)	(0.125)	(0.159)
Observations	3,894	1,663	989	911	1,099
R-squared	0.191	0.270	0.358	0.361	0.388

Notes: Robust standard errors in parentheses clustered at the járás (micro-region) level.

*** p<0.01, ** p<0.05, * p<0.1

A.7 Reduced form and 2SLS estimation results based on the specification of Lovász and Szabó-Morvai (2018)

Differences in empirical strategy compared to Lovász and Szabó-Morvai (2019)

Our empirical strategy also uses the birthdate of the smallest child as an instrument, but the treatment and control groups and the cut-off date differ from Lovász and Szabó-Morvai (2019) for two reasons: in contrast to the Census, mothers' outcomes are observed in all seasons of the year in LFS and the rules of admission to kindergarten has changed considerably since 2010. Before 2010, children could start kindergarten only after passing their third birthday. Lovász and Szabó Morvai (2019) used the informal cut-off rule for subsidized kindergartens, as follows: children who turned 3 before December 31 were able to enrol soon after their birthday, while those born after December 31 could only enrol in the following September. As a consequence, children, who were born in the autumn had a much higher probability to be in childcare than the children born between January- March within the months following shortly their third birthday. Against this background, Lovász and Szabó-Morvai (2019) define the treatment and control groups are defined such that they are observed when their children are the same age as follows. Mothers whose children were born between August 1 and December 31 constitute the treatment group, and they are observed in the quarter after their children turn 3 (between January 1 and March 31). The control group consists of mothers with child birthdates between January 1 and May 31, and observation dates the quarter after they turn 3 (June 1 to August 31).

From 2010, children who have reached the age of 2.5 years can also be admitted to a kindergarten, but this is subject to consideration by the kindergarten and depends on the available excess capacity. It implies that children who turn three by the end of February might be admitted to kindergarten in September of the previous year. However, the end-December remained the main important informal threshold in enrolment decisions, in the sense that kindergartens usually admit children at the beginning of a calendar year, in September who turns three by the end of that calendar year. Reflecting the change in admission rules, in our specification, mothers' outcomes are observed still before their child turns three and both the treatment and the control groups are observed in autumn.

This sampling design of Lovász and Szabó-Morvai (2019) ensures that child age and therefore the effect of parental leave and separation preferences will be the same on average in the two groups, in our specification, the child age slightly differs in the control and the treatment groups but the outcomes observed at the same period of the year, hence seasonality is not a concern. To address the potential difference between the control and treatment groups in terms of separation

preferences, we run the same estimation by shifting the treatment and control groups by +/- 6 months by the age of the youngest child.

The specification of Lovász and Szabó-Morvai (2019) does not show a significant effect of childcare availability on the employment of the mothers for any sample periods after 2005. The weakening estimated impact probably can be explained by the changing enrolment rules.

Table A.8: Reduced form and 2SLS results using Szabó-Morvai and Lovász's (2019) specification

Outcome:	Maternal employment						
Specification	Reduced form				2SLS estimation		
Cut-off	1st of January				1st of January		
Time period	1998-2011	2005-2019	2010-2015	2012-2019	2005-2019	2010-2015	2012-2019
T*m	0.090**	0.020	0.025	0.019			
	(0.027)	(0.027)	(0.037)	(0.034)			
C*m					0.028	0.036	0.026
					(0.038)	(0.050)	(0.046)
Constant	0.563*	0.406	1.096**	1.079*			
	(0.267)	(0.306)	(0.372)	(0.444)			
Observations	10041	6336	3328	4064	6330	3312	4043
R-squared	0.229	0.296	0.361	0.296	0.196	0.229	0.173

Notes: Robust standard errors in parentheses clustered at the catchment area level. All controls (individual, family and regional) included. Seasonally-corrected specifications (reduced form and 2SLS) with a 3-month time window.

*** p<0.01, ** p<0.05, * p<0.1