# Does formal child care availability for 0-3 year olds boost mothers' employment rate ? Panel data based evidence from Belgium

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## Abstract

In 2003, a new multi-annual program aimed at increasing the availability of formal child care for 0-3 year old children was launched in Wallonia, the French-speaking part of Belgium. This paper is interested in evaluating if this increased availability of formal child care resulted in a higher employment rate for women with at least one child under 3. To this end, we use a difference-in-differences approach based on municipality-level panel data, taking advantage of the fact that the increase in availability of formal child care differed greatly across municipalities. We find that the raise in child care availability significantly increased the maternal employment rate, but to a lesser extent than expected, most likely because of a substantial crowding-out effect.

## Résumé

En 2003, un nouveau programme pluriannuel visant à accroître la disponibilité de places d'accueil formelles pour les enfants de moins de 3 ans a été lancé en Wallonie, la région francophone de la Belgique. L'objet de ce papier est d'évaluer si cet accroissement de la disponibilité de places d'accueil a ou non entrainé une augmentation du taux d'emploi des mères ayant au moins un enfant agé de moins de 3 ans. Pour ce faire, nous utilisons une approche en double différence basée sur des données en panel au niveau municipal, tirant parti du fait que l'augmentation de la disponibilité de places d'accueil a grandement varié d'une municipalité à l'autre. Nos résultats montrent que l'accroissement de la disponibilité de places d'accueil a significativement augmenté le taux d'emploi des mères, mais de façon moins importante qu'espéré, vraisemblablement à cause d'un effet d'éviction substantiel.

**Keywords** : Child care, maternal employment, difference-in-differences, random trend model.
**JEL. classification** : J13, J21.

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# 1. Introduction

Despite significant progress in female labour force participation over the past decades, substantial gender differences remain. As demonstrated by Angrist and Evans (1998), and Moschion (2009) based on French data, one of the explanations for the lower female employment rate is the presence of children. In Belgium, according to the OECD family database 2014, the gender gap in employment rate is 10.9 percentage points. While the difference between the female employment rate and the maternal one is rather small compared to other OECD countries (73.8% and 70.6%, respectively), the maternal employment rate decreases drastically in the presence of children under 3 (62.1%).

Many factors such as education, labour market conditions or cultural idiosyncrasies influence maternal employment. Besides these factors, public policies are also important : tax policy, child benefits, paid leave, child care, etc. As a matter of fact, from studies based on international comparisons across European or OECD countries (see for example Gornick *et al.* (1997, 1998), Jaumotte (2003), Stadelmann-Steffen (2008), De Henau *et al.* (2010)), a consensus seems to emerge on the crucial role of family policies, and more specifically on the role played by the provision of formal child care.

In accordance with this state of the art, in 2002 the European Union recommended “to provide child care by 2010 to at least 90% of children between 3 years old and the mandatory school age, and to at least 33% of children under 3 years of age”. Following this recommendation, from 2003 onwards, the public agency in charge of formal child care in Wallonia – the French speaking part of Belgium – strongly increased the availability of formal child care places for children under 3. Nothing basically changed for children between 3 and the mandatory school age since, in Belgium, more than 90% of those children already attend preschool.

The purpose of this paper is to evaluate if this increased availability of formal child care in Wallonia resulted in a higher employment rate for women with at least one child under 3 years old. To this end, we use a difference-in-differences approach based on municipality-level panel data, taking advantage of the fact that the increased availability of formal child care differed greatly across municipalities.

This paper contributes to the literature investigating the effect of the availability and/or affordability of formal child care on maternal employment, based on the analysis of policy changes that create time and/or regional variations in the access to and/or price of child care (see among others Gelbach (2002), Schlosser (2005), Berlinski and Galiani (2007), Baker *et al.* (2008), Lefebvre and Merrigan (2008), Lundin *et al.* (2008), Cascio (2009), Fitzpatrick (2010, 2012), Goux and Maurin (2010), Simonsen (2010), Havnes and Mogstad (2011), Hardoy and Schøne (2013), Bauernschuster and Schlotter (2015), Bettendorf *et al.* (2015), Givord and Marbot (2015), Haeck *et al.* (2015), Lovász and Szabó-Morvai (2015), Nollenberger and Rodrigez-Planas (2015)). This literature, which mostly relies on difference-in-differences methods, has so far produced mixed results : some studies have found little effect (e.g., Havnes and Mogstad (2011)), while others have found relatively large effects (e.g., Bauernschuster and Schlotter (2015)). As outlined in Cascio *et al.* (2015), this is not really surprising as both the institutional backgrounds and the details of the policy changes analyzed in the different studies are very heterogeneous.

Most of the above mentioned papers deal with child care for children between 3 and 5 years old. Along with Goux and Maurin (2010), Simonsen (2010), Hardoy and Schøne (2013), Givord and Marbot (2015), Lovász and Szabó-Morvai (2015), and Nollenberger and Rodrigez-Planas (2015), this paper is one of the few available studies in this literature focusing specifically on formal child care for children under 3 years old. As far as we know, besides Vanleenhove (2013) who relies on simulations from an estimated structural labour supply model, it is the first policy evaluation study on the basis of Belgian data of the effect of child care on maternal employment.

In a nutshell, we find that the raise in child care availability significantly increased the maternal employment rate, but to a lesser extent than expected, most likely because of a substantial crowding-out effect: newly available formal child care seems to some extent to simply crowd out informal child care provided by other members of the family, friends or in the shadow economy.

The rest of the paper is organized as follows. Section 2 describes the institutional background of formal child care in Wallonia and the policy change that occurred in 2003. Section 3 discusses our empirical strategy. Section 4 describes the data. The results are presented in Section 5. Finally, Section 6 concludes.

# 2. Formal child care in Wallonia

## 2.1. Institutional background

In Wallonia,[1] the responsibility for formal child care falls under the auspices of ONE (Office de la Naissance et de l'Enfance, i.e., Birth and Children's Office), a public agency.

Formal child care services can be grouped into two broad categories: family care, provided by childminders at their private home, and collective care, mainly provided by day care centers.

Regardless of its type, each formal child care service has to be authorized by ONE. In order to be authorized, the care providers have to fulfill a number of conditions concerning dimensions of rooms and equipments, the child staff ratio, staff level of education and continuous training, medical follow-up for children, etc. The exact conditions depend on the type of the service. For example, the child staff ratio may at most be 4 to 1 for family care services, and 7 to 1 in collective care services (in this latter case, other adults such as nurses or social assistants may also be required).

The majority of child care services are subsidized by ONE. In order to be subsidized, services need to be accredited. This implies that besides being authorized, the child care service has to fulfill additional conditions. These conditions are mainly twofold: (a) in order to be accessible to working parents, the service must be opened at least 10 hours a day and 220 days a year, and (b) in order to guarantee finan-

[1] with the exception of nine municipalities belonging to its small German-speaking Community (about 2% of the Walloon population), which has its own separate public agency and will be ignored hereafter.

cial affordability, the day fee paid by parents must follow an income-dependent grid defined by ONE (the higher the household's income, the higher the daily fee).[2]

Subsidies by ONE are not the only source of public funding of formal child care services. Child care services may also (jointly or separately) benefit from employment and infrastructure related subsidies granted by the Walloon government and designed to help the non-profit sector, as well as from municipal grants. Only a small minority of child care services are managed without any public funding support.

On average, parental fees are lower in child care services subsidized by ONE than in non-subsidized ones, but it is worth noting that some of the non-subsidized services also apply the income-dependent grid defined by ONE.

## 2.2. Policy change

As already outlined, the European Union recommended in 2002 "to provide child care by 2010 to at least 90% of children between 3 years old and the mandatory school age, and to at least 33% of children under 3 years of age".

ONE monitors the availability of child care services by computing a "coverage rate" defined as the ratio of the number of available child care places to the number of children aged between 0 and 2.5 years old. The rationale for considering 0-2.5 year old children in the denominator is to approximate the number of children who potentially need child care services. It rests on the fact that, in Belgium, children usually enter child care services between 0 and 6 months and (pre-) school between 2.5 and 3 years.

According to this ONE indicator, in 2003, 20,933 places were available in Wallonia for 93,524 children, which represented a coverage rate of 22.4 %. It thus appeared that the fulfillment of the European Union objective required the creation of about 10,000 places, without taking into account a potential increase in birth rates.

In order to increase the number of available child care places, additional budgetary means were allocated to ONE. Thanks to these additional resources, in 2003 ONE launched a multi-annual program based on calls for projects. Projects were mainly submitted by various non-profit organizations and municipalities. They could consist of extensions of already existing child care services or the creation of new ones, with or without an infrastructure component. Projects were selected on the basis of a number of indicators at the municipality level, in order to promote a more universal access as well as to positively discriminate poor municipalities. To promote a more universal access, higher priority was given to projects in municipalities where the existing coverage rate (subsidized and non-subsidized by ONE) was lower than average, and where the birth rate and 18-45 female employment rate were higher than average. On the other hand, to positively discriminate poor municipalities, higher priority was given to projects in municipalities where the median income was lower than average, and the (overall) unemployment rate and proportion of low educated women were higher than average. The trade-off between these potentially conflicting objectives was done using a simple points system : for each indicator, the projects were classified by decile, on a scale ranging from 1 (lowest priority) to 10

[2] For more details, see Meulders and O'Dorchai (2008).

(highest priority), the final score of a project being obtained by summing its scores for the different indicators. At the same time, the Walloon government increased its employment and infrastructure related subsidies to the sector.

Overall, this multi-annual program has led to the opening of many new places. In 2010, some of the selected projects were not effective yet but nevertheless the target in terms of places was almost reached : 29,178 places were available in Wallonia, which represented 8,245 new places (+39.4 %) compared with the 20,933 places already available in 2003.

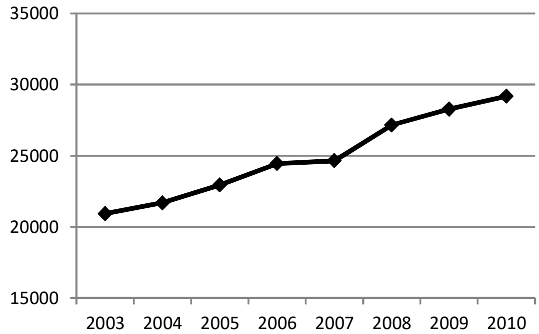



Figure 1 : Number of child care places in Wallonia

Due to an increase in birth rates over the same period, it led to a somewhat lower increase (+30%) in the coverage rate, which reached 29.2% in 2010, compared with 22.4% in 2003.

Whereas the increase in formal child care availability has been substantial in Wallonia, it has not been geographically homogeneous : some municipalities experienced a much larger increase in their coverage rate than the average (up to +136 %), while a few municipalities actually saw their coverage rate slightly decrease (up to -1%).

It is worth noting that during the same period parental fees as defined by ONE's income-dependent grid did not change (besides usual indexation). The policy change thus focussed on an increase in the availability of child care, without changing its financial affordability. Its primary goal was basically to work against the severe rationing (through waiting lists, where places are allocated on a first-come, first-serve basis) of the demand for child care.

To conclude on this section, a last remark. In 2002, when the European Union recommended "to provide child care by 2010 to at least 90% of children between 3 years old and the mandatory school age, and to at least 33% of children under 3 years of age", no exact indicators were defined to guide the European member states in checking whether or not they complied with the objectives. Regarding children under 3, as outlined above, the ONE indicator was used, and it appeared that Wallonia did not reach the 33% target. A harmonized indicator was finally defined by the European Union in 2004. This indicator strongly differs from the ONE indicator.[3] As a matter of fact, when the ONE first computed the European

[3] The denominator includes all children aged between 0 and 3 years old (instead of between 0 and 2.5 years old). The numerator includes both child care and preschool (instead of child care only) and it counts the number of registered children (instead of the number of places) between 0 and 3 years

harmonized indicator for Wallonia in 2009, it turned out to be equal to 48.3 %, i.e., largely in excess of the European objective. There is little doubt this was already the case in 2003.[4] To reach the European target was not the only argument when it was decided to boost the availability of formal child care : a strong consensus already existed that the supply of child care services was (and still is) insufficient. Along with the availability of new budgets,[5] the European objective nevertheless acted as a catalyst. Be that as it may, retrospectively, from the European recommendation point of view, the fact is that there was actually no need to boost formal child care availability for children under 3 in Wallonia.

# 3. Empirical strategy

We are interested in evaluating the effect of the increased availability of formal child care for children under 3 in Wallonia on mothers' employment rate. To this end, we use a difference-in-differences approach based on municipality-level panel data, exploiting the fact that the increased availability of formal child care widely varied across municipalities.

## 3.1. Model

Let $y_{it}$ denote the employment rate of women with at least one child under 3 in municipality $i$ at period $t$. For the sake of the argument, suppose that only two years ($t = 1, 2$) are observed and that the policy change between these two years has been such that the coverage rate (number of child care places per child) increased in some municipalities, while others remained unaffected. In other words, at period $t = 2$, some municipalities got 'treated', while others did not. Under these circumstances, a standard difference-in-differences estimator of the effect of the increased availability in formal child care may be obtained as the fixed effects (FE) or the first difference (FD) estimator of $\delta$ in the panel data regression model :[6]

$$y_{it} = c_i + \gamma d2_t + \delta D_{it} + \varepsilon_{it}, \quad i = 1, ..., N;\; t = 1, 2, \tag{1}$$

where $c_i$ is a municipality-specific effect, $d2_t$ is a time dummy equal to one if $t = 2$ and zero otherwise, and $D_{it}$ is a binary variable such that $D_{i1} = 0$ for all municipalities, and in period $t = 2$, $D_{i2} = 1$ for the municipalities where the coverage rate increased and $D_{i2} = 0$ otherwise.

The standard panel data difference-in-differences regression model (1) can be easily modified to accommodate our case where several years are observed and the treatment – the availability of child care – is continuous rather than binary, as well

old. Both the inclusion of preschool (in Belgium, more than 90% of children attend preschool) and the counting of registered children (as children may attend child care less than full time, a place may, on average, be occupied by more than one child) make a strong difference.

[4] The exact figure is not available for Wallonia, but for Belgium as a whole, when it was first computed in 2005, it was equal to about 42 %.

[5] In 2000 - 2001, a reform which gradually transferred financial resources from the federal government to the local governments was voted.

[6] See Wooldridge (2010, chapter 10), or Cameron and Trivedi (2005, chapter 22). Angrist and Pischke (2009, chapter 5) is also a useful reference. Note that the equality of the FE and FD estimator holds because there are only two periods.

as to relax its overly restrictive common trend assumption which supposes that the maternal employment rate would have evolved through time in the same way in all municipalities if there was no change in the availability of formal child care.

Allowing for more years of observation and for a continuous treatment is straightforward: it simply requires introducing a full set of time dummies and replacing the binary treatment with a variable measuring the intensity of the treatment, i.e., $z_{it}$ the coverage rate in municipality $i$ at period $t$. On the other hand, the common trend assumption can be relaxed first by allowing the time trend to be different across (a number $S$ of) sub-regions, in our case across provinces.[7] It can be further relaxed by allowing for municipality-specific time trend effects, finally yielding the so-called random trend model:[8]

$$y_{it} = c_i + g_i t + \sum_{s=1}^{S} ds_i(\gamma_{3s} d3_t + ... + \gamma_{Ts} dT_t) + \delta z_{it} + \varepsilon_{it}, \qquad \begin{array}{l} i = 1, ..., N, \\ t = 1, ..., T, \end{array} \tag{2}$$

where $g_i t$ is a municipality-specific time trend, $ds_i$ is a dummy equal to one if municipality $i$ belongs to sub-region $s$ and zero otherwise, and $dT_t$ is a time dummy equal to one if $t = T$ and zero otherwise.

The modified model (2) may be interpreted as a difference-in-differences model where each unit – the municipalities – is no longer either treated or not treated, but is put somewhere on a continuum of treatment, possibly including no (constant intensity of) treatment for some or all periods. The joint introduction of sub-region/time dummies and municipality-specific time trends allows the evolution over time of the maternal employment rate to fully differ across sub-regions, as well as to partially differ across municipalities within each sub-region. The sub-region/time dummies account for possibly different economic conditions across sub-regions. Along with the municipality-specific effects $c_i$, the municipality-specific time trends primarily account for differences in the composition of the population across municipalities, as such differences may not only imply differences in the level but also in the trend of employment rates.

We saw above that the increase in the availability of formal child care in Wallonia resulted from the adoption of a multi-annual program, which called for projects that were selected on the basis of indicators at the municipality level. Because the selection of projects was based on the municipality's characteristics, in model (2), the coverage rate $z_{it}$ can be expected to be correlated with both $c_i$ and $g_i$. On the other hand, because it resulted from the adoption of a multi-annual program, it may reasonably be assumed that the coverage rate $z_{it}$ is not systematically related to other factors than those captured by $(c_i, g_i)$ that influence the maternal employment rate $y_{it}$ (and that are left in $\varepsilon_{it}$). This is because in the context of a multi-annual program, the actual timing of the availability of the new child care places is essentially the result of an administrative process.[9] Under this assumption, the coverage rate $z_{it}$ may be considered as exogenous conditional on $(c_i, g_i)$ and model (2) may be viewed as a causal model.

[7] Wallonia is composed of five provinces (see Figure 3 in Section 4 below).

[8] See Wooldridge (2010, chapter 11).

[9] It depends on the pace at which the budgets are available, the time needed to extend or build new infrastructures, etc.

## 3.2. Estimation

Following Wooldridge (2010, chapter 11), let us rewrite model (2) as:

$$y_{it} = w_t a_i + x_{it}\beta + \varepsilon_{it}, \quad i = 1, ..., N;\ t = 1, ..., T, \tag{3}$$

where $w_t = (1, t)'$, $a_i = (c_i, g_i)$, $x_{it}$ is a row vector including the coverage rate $z_{it}$ and all the sub-region/time dummies, and $\beta$ a column vector of parameters composed of $\delta$ and all the $\gamma_{ts}$. Stacking the $T$ observations of each municipality, we can write:

$$Y_i = W a_i + X_i\beta + \varepsilon_i, \quad i = 1, ..., N. \tag{4}$$

We can get rid of the unobserved (and possibly correlated with $X_i$) municipality-specific effect and time trend $a_i$ by premultiplying equation (4) by the $T \times T$ transformation matrix $M = I_T - W(W'W)^{-1}W'$. This yields the transformed model:

$$\ddot{Y}_i = \ddot{X}_i\beta + \ddot{\varepsilon}_i, \quad i = 1, ..., N, \tag{5}$$

where $\ddot{Y}_i = MY_i$, $\ddot{X}_i = MX_i$ and $\ddot{\varepsilon}_i = M\varepsilon_i$. The transformation matrix $M$ is just a generalization of the usual fixed effects (or within) transformation.[10]

Under the strict exogeneity assumption $E(\varepsilon_i|X_i, a_i) = 0$, $i = 1, ..., N$, model (5) can be consistently estimated by pooled OLS ($N \to \infty$, $T$ fixed). This estimator is just a generalization of the usual fixed effects (FE) estimator and its asymptotic variance-covariance matrix may likewise be estimated using a robust estimator.

Following the approach underlying the fixed effects generalized least squares (FEGLS) estimator considered in Wooldridge (2010, chapter 10), a generally more efficient estimator may be obtained by taking into account – at least approximately – the second order moments of $\varepsilon_i$. In the present case, the idiosyncratic error $\varepsilon_{it}$ may be expected to be both serially correlated and heteroscedastic, with a variance inversely proportional to the number $n_i$ of women with at least one child under 3 in municipality $i$.[11] This suggests considering for the variance-covariance of $\varepsilon_i$:

$$V(\varepsilon_i|X_i, a_i) = h_i\Lambda, \quad i = 1, ..., N, \tag{6}$$

where the $h_i$ are known positive constants equal to $1/n_i$ and $\Lambda$ is an unrestricted $T \times T$ positive definite matrix. This specification is not assumed to be exact. It is simply intended to capture the most salient features of the second order moments of $\varepsilon_i$, in the hope of getting, through GLS, a more efficient estimator.

Under assumption (6), we have:

$$V(\ddot{\varepsilon}_i|\ddot{X}_i) = E(\ddot{\varepsilon}_i\ddot{\varepsilon}_i') = h_i M\Lambda M, \quad i = 1, ..., N, \tag{7}$$

which has rank equal to $T - 2$. The deficient rank of (7) makes the usual approach to GLS inapplicable. As in the case of Wooldridge's FEGLS estimator, the easiest

[10] Instead of considering data in deviation from their individual-specific mean, we thus consider data in deviation from their individual-specific level and trend.

[11] The observed maternal employment rate $y_{it}$ may be viewed as a sample frequency (the proportion of working mothers) estimated at the municipality level, whose variance is accordingly inversely proportional to the size of the population of mothers of the municipality. In our dataset, municipality size varies by a factor as large as 75. It is thus important to take it into account.

way to bypass this problem is to drop two of the time periods from the analysis, say the first two time periods.

To avoid introducing new notations, we now let $\ddot{Y}_i$, $\ddot{X}_i$ and $\ddot{\varepsilon}_i$ denote, respectively, the $(T-2)\times 1$ vector, the $(T-2)\times K$ matrix and the $(T-2)\times 1$ vector obtained after dropping the first two time periods from the transformed model (5). Accordingly let $V(\ddot{\varepsilon}_i|\ddot{X}_i)$ be written as $V(\ddot{\varepsilon}_i|\ddot{X}_i) = E(\ddot{\varepsilon}_i\ddot{\varepsilon}_i') = h_i\Omega$, where $\Omega$ is an unrestricted $(T-2)\times(T-2)$ positive definite matrix.[12] A generalized version of Wooldridge's FEGLS estimator is given by:

$$\hat{\beta} = \left(\sum_{i=1}^{N} \frac{1}{h_i} \ddot{X}_i' \hat{\Omega}^{-1} \ddot{X}_i\right)^{-1} \left(\sum_{i=1}^{N} \frac{1}{h_i} \ddot{X}_i' \hat{\Omega}^{-1} \ddot{Y}_i\right), \tag{8}$$

where $\hat{\Omega}$ is a consistent estimator of $\Omega$. Such an estimator is given by:

$$\hat{\Omega} = \frac{1}{N} \sum_{i=1}^{N} \frac{1}{h_i} \widetilde{\ddot{\varepsilon}}_i \widetilde{\ddot{\varepsilon}}_i', \tag{9}$$

where $\widetilde{\ddot{\varepsilon}}_i$ denote the $(T-2)\times 1$ vector of residuals $\widetilde{\ddot{\varepsilon}}_i = \ddot{Y}_i - \ddot{X}_i\tilde{\beta}$, $i = 1, ..., N$, calculated using any first step consistent estimator $\tilde{\beta}$ of $\beta$, for example the weighted pooled OLS estimator obtained by replacing $\hat{\Omega}$ by an identity matrix in (8). A fully robust (to misspecification of (6)) variance-covariance matrix estimator for the generalized FEGLS estimator $\hat{\beta}$ is given by:

$$\widehat{V}(\hat{\beta}) = \left(\sum_{i=1}^{N} \frac{1}{h_i} \ddot{X}_i' \hat{\Omega}^{-1} \ddot{X}_i\right)^{-1} \left(\sum_{i=1}^{N} \frac{1}{h_i^2} \ddot{X}_i' \hat{\Omega}^{-1} \widehat{\ddot{\varepsilon}}_i \widehat{\ddot{\varepsilon}}_i' \hat{\Omega}^{-1} \ddot{X}_i\right) \left(\sum_{i=1}^{N} \frac{1}{h_i} \ddot{X}_i' \hat{\Omega}^{-1} \ddot{X}_i\right)^{-1},$$

where $\widehat{\ddot{\varepsilon}}_i$ stands for the $(T-2)\times 1$ vector of generalized FEGLS residuals $\widehat{\ddot{\varepsilon}}_i = \ddot{Y}_i - \ddot{X}_i\hat{\beta}$, $i = 1, ..., N$.

# 4. Data

This study relies on data from administrative sources. The municipality-level employment rates come from the Crossroads Bank for Social Security (CBSS), a public agency gathering administrative data on social security (employment, welfare, invalidity, retirement, etc.) for all Belgian residents. With the CBSS database, it is possible to compute, for each Walloon municipality, the employment rate of 18-49 year old women with at least one child under 3. The employment rate for year $t$ refers to the employment situation at the end of the year (December, 31).

Employment data were originally collected for the 2003-2009 period. However, it turned out that there was a break in the series in 2004, due to a definitional change. Our analysis is therefore restrained to the 2005-2009 period. Due to privacy laws, we had to regroup the smallest municipalities with their neighboring municipality. Also, one municipality has been excluded from the analysis because it is geographically isolated (no joint frontier with any other Walloon municipality). As a result, out of

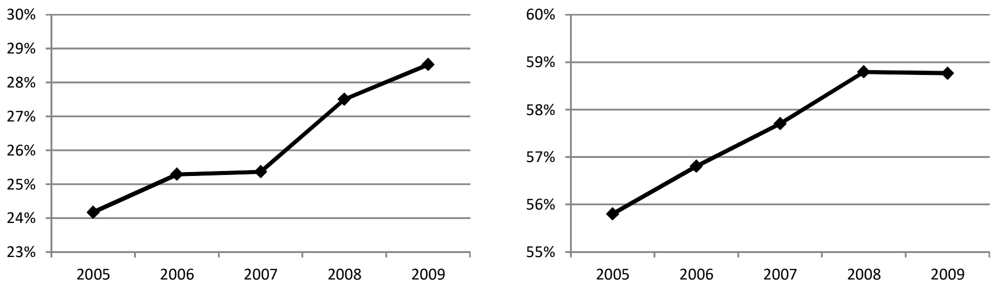
[12] Since $\Lambda$ is assumed unrestricted, $\Omega$ is also unrestricted.

253 municipalities, we finally ended up with observations on 235 distinct entities, which we will hereafter continue to call municipalities for simplicity.

The municipality-level coverage rates come from ONE and Statistics Belgium, the public agency that manages the National Register database. For each Walloon municipality and each year from 2003 to 2010, ONE provided the number of authorized child care places and Statistics Belgium the number of children aged between 0 to 2.5 years. The ratio defines the standard ONE coverage rate intended to measure the availability of formal child care.[13] As for the employment rate, data for year $t$ refers to the situation at the end of the year (December, 31).

It is rather unrealistic to assume that when parents are looking for child care services, they only consider services located in their municipality of residence. It is more plausible that they look for available services in a wider area. Accordingly, hereafter in our analysis, the coverage rate of a municipality is defined as the coverage rate (number of child care places per child) over an enlarged area, comprising the considered municipality and its surrounding (contiguous) municipalities.

Figure 2 depicts the aggregate child care coverage rate and employment rate of women with at least one child under age 3 for Wallonia as a whole over our 2005-2009 period of analysis.




Coverage rate

Employment rate

Figure 2 : Child care coverage rate and employment rate of women with at least one child under age 3 in Wallonia

From 2005 to 2009, the coverage rate increased from 24.17% to 28.53%. The employment rate of women with at least one child under age 3 increased from 55.80% to 58.77%, with virtually no change between 2008 and 2009, undoubtedly as a result of the 2008 financial crisis.

The aggregate child care coverage rate and women's employment rate for Wallonia as a whole hide huge differences across municipalities. According to Table 1, the child care coverage rate varied across municipalities from 12,48% to 60,53% in 2005, and from 15.65% to 67.09% in 2009. Likewise, women's employment rate varied across municipalities from 20.42% to 83.87% in 2005, and from 24.00% to 85.12% in 2009.

[13] As a reminder, the rationale for considering 0-2.5 year old children in the denominator is to approximate the actual number of children who potentially need child care services.

Table 1 : Child care coverage rate and employment rate of women with at least one child under age 3 across municipalities

| Variable | Min. | Quart. 1 | Median | Quart. 3 | Max. |
|---|---|---|---|---|---|
| Coverage rate in 2005 | 12.48 | 20.33 | 24.62 | 28.38 | 60.53 |
| in 2009 | 15.65 | 23.94 | 30.64 | 34.78 | 67.09 |
| Employment rate in 2005 | 20.42 | 52.61 | 64.56 | 71.41 | 83.87 |
| in 2009 | 24.00 | 57.88 | 67.74 | 75.66 | 85.12 |

Notes : The coverage and employment rates are in percentage. Observations from 235 municipalities.

Also, the increase in the aggregate child care coverage rate between 2005 and 2009 for Wallonia as a whole – from 24.17% to 28.53%, that is to say +18.03% – hides large differences across municipalities. As illustrated by Figure 3, the relative increase in the coverage rate between 2005 and 2009 varied across municipalities from -4.09% to +92.82%. In other words, the change in availability of formal child care differs across municipalities from a slight decrease to an almost doubling over the 2005-2009 period.

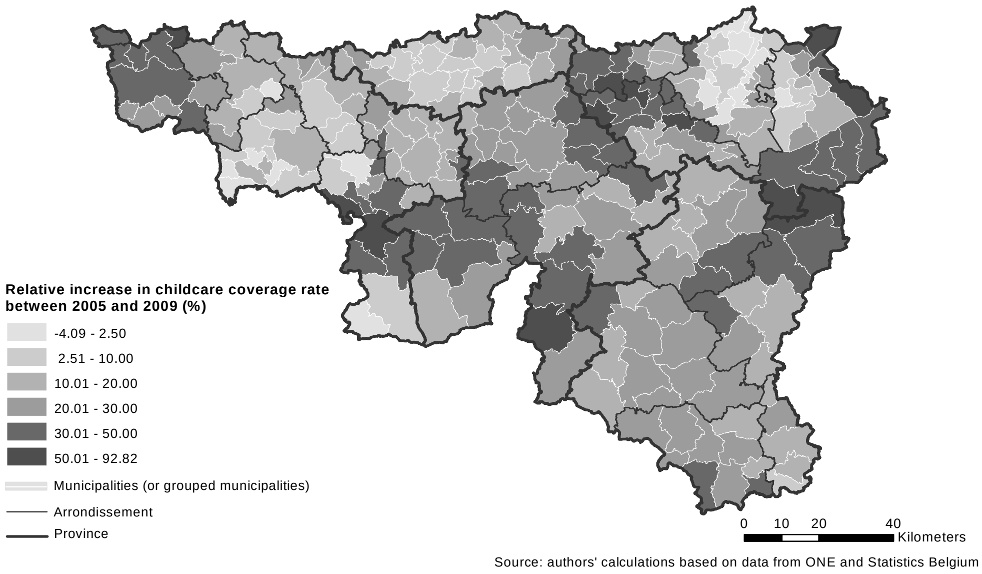



Figure 3 : Relative increase in child care coverage rate between 2005 and 2009 across municipalities

# 5. Results

## 5.1. Benchmark results

Table 2 reports the results of the generalized FEGLS estimation of model (2) for 18-49 year old women with at least one child under 3. For comparison purposes, the results of the estimation of the same model for 18-49 year old men with at least one child under 3, as well as for 18-49 year old men and women without children, are also reported.[14] For reasons of conciseness, only the estimated parameter of interest $\delta$ and its standard error are reported.

Table 2 : Benchmark results

| Variable | Women with at least one child under age 3 | Men with at least one child under age 3 | Women without children | Men without children |
|---|---|---|---|---|
| Coverage rate | 0.176*** | 0.019 | 0.005 | 0.023 |
| | (0.065) | (0.049) | (0.057) | (0.051) |

Notes : Generalized FEGLS estimates. 235 municipalities observed over 2005-2009. Robust standard errors in parentheses. Significance level : * = 10%, ** = 5% and *** = 1%.

As it may be seen, the availability of formal child care as measured by the coverage rate is found to have a highly significant and positive effect on the employment rate of mothers with at least one child under age 3. In contrast, the availability of formal child care appears to have no significant effect on the employment rate of fathers with at least one child under age 3, and likewise no significant effect on the employment rate of men and women without children.

The absence of effect for fathers with at least one child under age 3 is not surprising : it simply corroborates the fact that there are still large gender differences with regard to child care. The absence of effect for women (and men) without children was expected. This result clearly supports our empirical strategy, in particular regarding the question of whether the inclusion of municipality-specific (level and) time trend effects is enough to purge the possible endogeneity of the coverage rate resulting from the selection of projects by ONE based on municipal characteristics. If it was not enough, the observed effect for women with children might to some extent be spurious, but should likewise be observed for women without children. It does not appear to be the case.

According to Table 2, an increase of one percentage point in the coverage rate yields an increase of 0.176 percentage points in the employment rate of women with at least one child under 3. Because the ratio between the number of children aged 0 to 2.5 and the number of women with at least one child under age 3 is about equal to one, this means that when 100 new child care places are opened, about 18 additional women take up paid work. For comparison purposes, similar estimates of the percentage point increase in the maternal employment rate per percentage point increase in the child care coverage rate reported in the literature range from about 0.05 (Havnes and Mogstad (2011)) to about 0.35 (Bauernschuster and Schlotter (2015)). Among the studies focusing specifically on formal child care for children under 3 years old, both Bettendorf *et al.* (2015) and Nollenberger and Rodrigez-Planas (2015) report estimates around 0.19, i.e., comparable to our finding. However, Lovász and Szabó-Morvai (2015) found a much lower effect (around 0.08), while the results reported in Haeck *et al.* (2015) suggest a much higher effect (around 0.30).[15]

[14] The municipality-level employment rates for these three other categories were likewise obtained from the CBSS database.

[15] The papers of Bettendorf *et al.* (2015) and Haeck *et al.* (2015) do not specifically deal with child care for children under 3 years old, but do provide disaggregate results for this category.

Tables 3 and 4 present some specification tests and sensitivity analysis allowing to assess the robustness of our estimated effect. Table 3 reports variable addition type tests intended to check, on the one hand, the linearity of the effect, and on the other hand, the strict exogeneity – conditional on $(c_i, g_i)$ – of the coverage rate.

Table 3: Women with at least one child under age 3
Specification tests

| Variable | Benchmark model | Alternative specification (1) | (2) | (3) | (4) |
|---|---|---|---|---|---|
| Coverage rate | 0.176*** | 0.164** | 0.190*** | 0.184*** | 0.198** |
| | (0.065) | (0.071) | (0.065) | (0.071) | (0.078) |
| Squared coverage rate | – | 0.001 | – | – | 0.001 |
| | | (0.003) | | | (0.003) |
| Lag of coverage rate | – | – | 0.050 | – | 0.071 |
| | | | (0.065) | | (0.064) |
| Lead of coverage rate | – | – | – | 0.029 | 0.058 |
| | | | | (0.074) | (0.072) |

Notes: The coverage rate is centered at the full sample mean. Generalized FEGLS estimates. 235 municipalities observed over 2005-2009. Robust standard errors in parentheses. Significance level: * = 10%, ** = 5% and *** = 1%.

The linearity of the effect may readily be tested by adding the squared coverage rate to the model. As it may be seen, there is no sign of non linearity, and in particular no sign that the marginal effect of the coverage rate decreases as the coverage rate itself increases, at least over the observed range of the coverage rate (reported in Table 1). In other words, an increase of the availability in formal child care has not only an effect when it is low, but also at higher absolute levels.

The strict exogeneity – conditional on $(c_i, g_i)$ – of the coverage rate may be checked by adding one lead and one lag of the coverage rate to the model.[16] The lag of the coverage rate allows to capture a possible lagged effect of child care availability, while the lead of the coverage rate allows to check for the presence of any anticipation effect or any feedback effect of the current employment rate on the future coverage rate. The presence of any of these effects would jeopardize the strict exogeneity – conditional on $(c_i, g_i)$ – assumption on which relies the validity of our generalized FEGLS estimator.[17] As it may be seen, there is no sign of any of these effects.

Table 4 shows the sensitivity of the estimated effect of the coverage rate to some variations in our benchmark model. Table 4 first reports what happens when (1) no municipality-specific time trend is included in the model (i.e., no $g_i t$ term is included in the model), (2) the aggregate time trend of the model is not allowed to fully differ across provinces (i.e., the full set of sub-region/time dummies is replaced by a set of time dummies only), and (3) interaction terms between the aggregate time trend and the municipality indicators on which were based the selection of projects by

[16] Thanks to the fact that the coverage rates were available for both 2004 and 2010, this could be done without losing any time period for estimation.

[17] For a discussion on strict exogeneity in panel data models with unobserved individual effects, see Wooldridge (2010, chapter 10).

ONE[18] are added to the model. It appears that allowing for municipality-specific time trends is very important: without such municipality-specific time trends, the estimated effect is almost divided by two (from 0.176 to 0.096). On the other hand, not allowing for fully different aggregate time trends across provinces has a much lower impact on the estimated effect, while adding interaction terms between the aggregate time trend and the municipality indicators on which were based the selection of projects by ONE leaves virtually unchanged the estimated effect.

Table 4: Women with at least one child under age 3
Sensitivity analysis

| | Coverage rate | |
|---|---|---|
| | Parameter | Std. error |
| Benchmark model | 0.176*** | 0.065 |
| Variation from benchmark model: | | |
| (1) No municipality-specific time trend | 0.096* | 0.056 |
| (2) No different aggregate trends across provinces | 0.139** | 0.058 |
| (3) Added interactions between aggregate trend and project's selection variables | 0.180*** | 0.062 |
| (4) Coverage rate defined without surrounding municipalities | 0.070*** | 0.026 |
| (5) Coverage rate defined at the level of arrondissements | 0.203** | 0.102 |
| (6) Municipalities with "extreme" coverage rate excluded | 0.149** | 0.072 |
| (7) Municipalities with "extreme" employment rate excluded | 0.191*** | 0.069 |

Notes: Generalized FEGLS estimates. 235 municipalities observed over 2005-2009, except (6) and (7) where 24 municipalities are excluded. Robust standard errors. Significance level: * = 10%, ** = 5% and *** = 1%.

As the finding of no effect for women without children, these results are reassuring regarding the question of whether our estimated effect might to some extent be spurious. One might face a spurious effect if, on the one hand, the selection of projects by ONE had been such that the coverage rate increased more in municipalities where the employment rate would have also increased more in absence of child care expansion, and on the other hand, as already mentioned, the inclusion of municipality-specific (level and) time trend effects in the model was not enough to control for this. The fact that the estimated effect is (drastically) reduced without municipality-specific time trends however suggests a negative correlation between the coverage rate $z_{it}$ and the (unobserved) municipality-specific time trends $g_i t$, i.e., that the selection of projects by ONE has on the contrary actually been such that the coverage rate increased more in municipalities where the employment rate would

[18] According to the selection process described in Section 2.2, the considered indicators are the existing coverage rate (subsidized and non-subsidized by ONE), the birth rate, the 18-45 female employment rate, the median income, the (overall) unemployment rate and the proportion of low educated women (at the time of selection). Data come from Statistics Belgium, WalStat, and the 2001 Census for the proportion of low educated women.

have increased less in absence of child care expansion. The fact that the addition of interaction terms between the aggregate time trend and the municipality indicators on which was based the selection of projects by ONE leaves virtually unchanged our estimated effect suggests that the inclusion of municipality-specific (level and) time trend effects does indeed already make a good job at controlling for the possible endogeneity of the coverage rate resulting from the selection process of projects.

Table 4 further shows how the estimated effect changes when (4) the coverage rate is not defined over an enlarged area comprising the considered municipality and its surrounding (contiguous) municipalities, but at the level of the considered municipality only, and (5) the coverage rate is defined over an even larger area, at the level of arrondissements, each made up of several municipalities (see the map in Figure 3). It may be seen that when only the municipality of residence is considered, the estimated effect is drastically reduced (from 0.176 to 0.070). This may be viewed as a consequence of the fact that the coverage rate defined at the level of the considered municipality only provides a rather bad measurement of the actual availability of formal child care : as we already argued, it is unrealistic to assume that when parents are looking for child care services, they only consider services located in their municipality of residence. Defining the coverage rate over an enlarged area is definitely a better – although still approximative – choice. Interestingly, it appears that the exact definition of this enlarged area is not critical : the estimated effect remains of the same magnitude regardless of whether the coverage rate is defined at the level of arrondissements or at the level of the considered municipalities plus their surrounding (contiguous) municipalities as in our benchmark model.

Table 4 finally reports what happens when (6) the municipalities with the 5% highest and 5% lowest coverage rates (in 2005) are excluded from the sample, and (7) the municipalities with the 5% highest and 5% lowest employment rates (in 2005) are likewise excluded from the sample. We saw in Section 4 that both the coverage rate and the employment rate tremendously vary across municipalities, exhibiting extreme minimum and maximum values (see Table 1). One might be concerned that the estimated effect is overly driven by these extreme values. However, this does not appear to be the case : the magnitude of the estimated effect is only marginally affected when the municipalities exhibiting these extreme values are excluded from the sample.

Our analysis implicitly assumes that the increased availability of formal child care had no side effect on mobility or fertility. If it actually had an impact on mobility or fertility, this might to some extent bias our estimated effect. As a matter of fact, some working women having or who wanted to have children might have been encouraged to move to municipalities where the coverage rate increased more. Likewise, women who want to combine work and family life might have been encouraged to have or to have more children in municipalities where the coverage rate increased more. In both cases, this would bias our estimated effect upwards. To shed light on these questions, Table 5 reports the results of the generalized FEGLS estimation of the same model as model (2), but where the dependent variable $y_{it}$ is respectively the net immigration rate and the birth rate of women aged 18 to 49 in municipality $i$ at time $t$.[19]

[19] The net immigration rate is defined as the difference between the number of entries and exits of

Table 5: Women aged 18 to 49
Mobility and fertility

| | Dependent variable | | | | | |
|---|---|---|---|---|---|---|
| Variable | Net immigration rate | | | Birth rate | | |
| Coverage rate | -0.041 | – | -0.044 | 0.025 | – | 0.028* |
| | (0.037) | | (0.037) | (0.016) | | (0.016) |
| Lag of coverage rate | – | -0.009 | -0.015 | – | 0.009 | 0.014 |
| | | (0.032) | (0.031) | | (0.013) | (0.013) |

Notes: Generalized FEGLS estimates. 235 municipalities observed over 2005-2009. Robust standard errors in parentheses. Significance level: * = 10%, ** = 5% and *** = 1%.

As it may be seen, we find no evidence of a side effect on mobility: neither the coverage rate nor its lag appears to have a significant effect on the net immigration rate of women aged 18 to 49. Also, there is only statistically fragile evidence of a possible side effect on fertility: the coverage rate is found to have a significant effect on the birth rate of women aged 18 to 49, but only at the 10% level and when its (insignificant) lag is jointly introduced in the model.

Overall, the results from Tables 3, 4 and 5 suggest that our estimated effect of the availability of formal child care (as measured by the coverage rate) on the employment rate of mothers with at least one child under age 3 is rather robust. However, with only about 18 additional women taking up paid work for 100 new child care places, the effect may seem somewhat disappointing.

This moderate effect may in part be explained by the fact that we consider the effect of child care availability on the maternal employment rate, an outcome variable which measures the proportion of mothers at work, but not the intensity of their work supply. As a matter of fact, when new child care places are opened, additional women may be induced to work, but women who already worked may likewise be induced to increase their working time, for example from part-time to full-time. These changes in working hours are ignored by the present study as it focuses on the employment rate.

Another partial explanation could be that our measurement of child care availability is not sufficiently accurate. As already outlined, defining the coverage rate over an enlarged area is definitely a better choice than at the level of the considered municipality only, but it is still approximative: parents may for example not only look for child care services located close to their place of residence, but also close to their place of work. This measurement error might to some extent bias downward our estimated effect.

The presumably most important explanation for the moderate estimated effect is that the increase in formal child care availability actually yielded some women who would otherwise have resorted to informal child care – i.e., child care provided by other members of the family, friends or in the shadow economy – to now resort to the newly available formal child care services. As suggested by the available literature

women during year $t$, divided by the number of women at the beginning of the year. Likewise, the birth rate is defined as the ratio between the number of births during year $t$ and the number of women at the beginning of the year. Data come from Statistics Belgium.

(see among others Baker *et al.* (2008), Cascio (2009), Havnes and Mogstad (2011), Bettendorf *et al.* (2015), Cascio *et al.* (2015), Givord and Marbot (2015), Haeck *et al.* (2015), and for Belgium Vanleenhove (2013)), this crowding-out effect is most likely to be substantial.

## 5.2. Extensions

So far, we did not make any distinction among formal child care services. A first distinction can be made regarding their location, either in the considered municipality or in its surrounding (contiguous) municipalities.[20] Another distinction can be made between family care, provided by childminders at their private home, and collective care, mainly provided by day care centers. Finally, a further distinction can be made between child care services that are subsidized by ONE, and those that are not. For Wallonia as a whole, the part of subsidized services was equal to 70.77% in 2005 and 72.30% in 2009, while the part of collective services was equal to 54.28% in 2005 and 54.90% in 2009.[21] The part of strictly local services (i.e., located in the considered municipality) was on average equal to 14.25% in 2005 and 14.62% in 2009.[22] These figures hide large differences across municipalities, both in level and over time.

Theoretically, a larger part of subsidized services should boost the maternal employment rate as a result of their better financial affordability.[23] It may also be argued that a larger part of collective services might likewise foster the maternal employment rate due to the fact that in principle collective services more easily satisfy "emergency" child care requests (for going to a job interview, participating in a training, or to respond positively to an unexpected job offer). A larger part of strictly local services might also be beneficial as it may allow to save time and travel costs. These conjectures may readily be checked by adding the part of subsidized services, collective services and strictly local services to our benchmark model.

As it may be seen from Table 6, neither the part of subsidized services nor the part of collective services nor the part of strictly local services appears to have a significant effect on the maternal employment rate. The apparent absence of effect of the part of subsidized services is most likely related to the fact that the part of subsidized services is already high in most municipalities (it is above 45% in 95% of municipalities). It may also partly be explained by the fact that some non-subsidized services also apply the income-dependent grid defined by ONE. The apparent absence of effect of the part of collective services and the part of strictly local services suggests that the exact mix between family and collective services, as well as between strictly local and surrounding services, is in practice of secondary importance.

[20] Recall that we defined the coverage rate over an enlarged area, comprising the considered municipality and its surrounding (contiguous) municipalities.

[21] The part of subsidized (resp. collective) services is defined as the number of places available in subsidized (resp. collective) services divided by the total number of available child care places.

[22] The part of strictly local services is defined as the number of places available in the considered municipality divided by the total number of places available in both the considered municipality and its surrounding (contiguous) municipalities.

[23] As a reminder, in subsidized services, the day care fee paid by parents must follow an income-dependent grid defined by ONE.

The increase in the availability of formal child care may not boost the employment of all mothers equally. In particular, it might be less effective for low-educated and/or single mothers, notably because of the existence of unemployment traps. A similar concern may be raised for mothers living in rural areas. Because the proportions of low-educated women and of single mothers vary across municipalities, the first conjecture may be assessed – admittedly in a somewhat crude way[24] – by adding to our benchmark model interaction terms between the coverage rate and dummy variables identifying respectively municipalities with a high proportion of low-educated women[25] and municipalities with a high proportion of single mothers.[26] The latter conjecture may readily be checked by further adding to the model an interaction term between the coverage rate and a dummy variable identifying rural municipalities.[27]

Table 6 : Women with at least one child under age 3
Extensions

| Variable | Benchmark model | Extension (1) | (2) | (3) |
|---|---|---|---|---|
| Coverage rate | 0.176*** | 0.262*** | 0.205** | 0.266*** |
| | (0.065) | (0.084) | (0.082) | (0.093) |
| Part of subsidized services | – | 0.021 | – | 0.008 |
| | | (0.044) | | (0.043) |
| Part of collective services | – | -0.064 | – | -0.055 |
| | | (0.040) | | (0.040) |
| Part of strictly local services | – | 0.044 | – | 0.044 |
| | | (0.047) | | (0.047) |
| Coverage rate × high proportion of low-educated women dummy | – | – | -0.261** | -0.238* |
| | | | (0.130) | (0.135) |
| Coverage rate × high proportion of single mothers dummy | – | – | -0.117 | -0.105 |
| | | | (0.154) | (0.159) |
| Coverage rate × rural municipality dummy | – | – | 0.298** | 0.283** |
| | | | (0.137) | (0.139) |

Notes : Generalized FEGLS estimates. 235 municipalities observed over 2005-2009. Robust standard errors in parentheses. Significance level : * = 10%, ** = 5% and *** = 1%.

As shown by Table 6, the effect of the coverage rate on the maternal employment rate is found to be significantly lower in municipalities with a high proportion of low-educated women. This suggests that only increasing the availability of formal

[24] A proper investigation of these questions would require individual data. The results reported here should therefore be viewed as merely tentative.

[25] These are defined as the municipalities where, according to the 2001 Census data, (1) the proportion of women with primary education is higher than 16% (this proportion varies from 5.37% to 27.18% across municipalities, with a median equal to 15.14%), and (2) the proportion of women with higher education is lower than 27% (this proportion varies from 13.79% to 52.37% across municipalities, with a median equal to 28.64%). These represent about 27% of our sample.

[26] These are defined as the municipalities where, according to our CBSS data, the average proportion of single mothers over the 2005-2009 period is higher than 15.5% (this average proportion varies from 1.44% to 31.86% across municipalities, with a median equal to 10.08%). These represent about 30% of our sample.

[27] These are defined as the municipalities classified in 2008 as “thinly populated areas” according to the Degree of Urbanisation (DEGURBA) classification of EUROSTAT. These represent about 28% of our sample.

child care is probably not enough to really boost the employment of low-educated mothers. On the other hand, the effect of the coverage rate is not found to be significantly lower in municipalities with a high proportion of single mothers, so that it cannot be argued this is also the case for single mothers. Finally, it appears that the effect of the coverage rate on the maternal employment rate is significantly higher in rural municipalities. The reasons underlying this somewhat unexpected result are unclear. A tentative explanation might be that there are less possibilities of informal child care arrangements in rural areas, so that the crowding out effect is less important.

## 5.3. Aggregate effect

For Wallonia as a whole, the employment rate of 18-49 year old women with at least one child under 3 increased from 55.80% in 2005 to 58.77% in 2009. Using our estimated results, the effect of the 2005-2009 change in available child care on the maternal employment rate may be evaluated for each municipality, then aggregated for Wallonia as a whole, allowing thereby to deduce what would have been the aggregate maternal employment rate in 2009 if child care availability remained at its 2005 level. Table 7 reports the result of this calculation, based on both the estimates of our benchmark model and of the extended model (3) of Table 6 to assess its robustness.

Table 7 : Women with at least one child under age 3
Aggregate effect of child care availability on employment rate

| | Benchmark model | Extended model |
|---|---|---|
| Employment rate in 2005 | 55.80 | |
| Employment rate in 2009 | 58.77 | |
| Effect of the 2005-2009 increase of child care availability on employment rate | +0.75<br>[+0.20 , +1.29] | +0.88<br>[+0.13 ; 1.63] |
| Hypothetical employment rate in 2009 with child care availability of 2005 | 58.02<br>[57.48 , 58.56] | 57.89<br>[57.13 , 58.64] |

Notes : The employment rates are in percentage. The extended model refers to the extended model (3) of Table 6. 95% confidence intervals between brackets.

As it may be seen, using our benchmark model, the effect of the 2005-2009 increase in child care availability on the aggregate employment rate of Walloon women with at least one child under age 3 is estimated to be equal to 0.75 percentage points. Accordingly, if child care availability remained at its 2005 level, it is estimated that in 2009 the aggregate maternal employment rate would have been equal to 58,02%, instead of 58.77%. This represents a loss of about 750 working mothers. Viewed in another way, about 25% of the increase in the maternal employment rate observed over the 2005-2009 period in Wallonia may be attributed to the increased availability in formal child care. Using the extended model leads to estimates of the same magnitude.

# 6. Conclusion

This paper was interested in evaluating the effect of the increased availability of formal child care for children under 3 in Wallonia on the employment rate of their mothers. To this end, we used a difference-in-differences approach based on municipality-level panel data, taking advantage of the fact that the increased availability in formal child care widely varied across municipalities. We found that an increase of one percentage point in the coverage rate yields an increase of about 0.18 percentage points in the employment rate of women with at least one child under 3, which means that when 100 new child care places are opened, about 18 additional women take up paid work. This moderate effect is most likely due to a substantial crowding-out effect.

These results suggest that no more than a moderate effect on maternal employment should be expected from a further similar increase in the availability of formal child care. This does not mean that such a policy is worthless. It must be stressed that supporting maternal employment is only one of the goals of public policies aimed at increasing the availability and/or affordability of formal child care. Other goals, which are perhaps even more important from the point of view of policy makers, are the improvement of children's cognitive and social development, as well as equity.

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