Local childcare availability and dual-earner fertility : variation in childcare coverage and birth hazards over place and time

# Local Childcare Availability and Dual-Earner Fertility

## Variation in Childcare Coverage and Birth Hazards over Place and Time

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The theoretically well-grounded hypothesis that the availability of formal childcare has a positive impact on childbearing in the developed world has been part of the population literature for a long time. Whereas the participation of women in the labour force created a tension between work and family life, the increasing availability of formal childcare in many developed countries is assumed to reconcile these two life domains due to lower opportunity costs and compatible mother and worker roles. However, previous empirical studies on the association between childcare availability and fertility exhibit ambiguous results and considerable variation in the methods applied. This study assesses the childcare–fertility hypothesis for Belgium, a consistently top-ranked country concerning formal childcare coverage that also exhibits considerable variation within the country. Using detailed longitudinal census and register data for the 2000s combined with childcare coverage rates for 588 municipalities and allowing for the endogenous nature of formal childcare and selective migration, our findings indicate clear and substantial positive effects of local formal childcare provision on birth hazards, especially when considering the transition to parenthood. In addition, this article quantifies the impact of local formal childcare availability on fertility at the aggregate level and shows that in the context of low and lowest-low fertility levels in the developed world, the continued extension of formal childcare services can be a fruitful tool to stimulate childbearing among dual-earner couples.

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## 1. Introduction

The second half of the twentieth century has been characterized by the rise of the dual earner model in many developed countries. Women's increased economic activity introduced a tension between work and family (Rindfuss, Guilkey, Morgan, Kravdal, & Guzzo, 2007). Sociologists stress the incompatibility between mother and worker roles, while economists highlight the increased opportunity costs of non-market activities such as childrearing tasks. Rising female labour force participation and declining fertility in the developed world since the 1960s suggest that childbearing is compromised to safeguard labour market participation. As a result, well-established fertility theories typically attributed the decline of fertility in the developed world to women's increased socio-economic position and individualistic attitudes geared towards career development rather than childrearing (Becker, 1981; Van de Kaa & Lesthaeghe, 1986).

In response to declining fertility levels and an increasing female labour supply, Western European governments have continuously extended family policies geared towards the reconciliation of work and family such as formal childcare (Rindfuss & Brewster, 1996; Thevenon, 2008). Despite indications of higher fertility for women with high socio-economic positions in contexts with extensive work-family reconciliation policies (Matysiak & Vignoli, 2008; Puur, Klesment, Rahnu, & Sakkeus, 2016; Wood & Neels, 2017; Wood, Neels, & Kil, 2014), empirical evidence on the effect of such policies remains inconclusive (Gauthier, 2007; Neyer & Andersson, 2008). This paper argues that the mixed empirical support for the theoretically well-grounded childcare-fertility hypothesis is, at least partially, related to the broad variation in methods applied. First, variation in formal childcare and fertility has been studied at different levels. Numerous cross-national comparisons included childcare indicators at the country level (Luci & Thevenon, 2012; Puur et al., 2016; Van Bavel & Rozanska-Putek, 2010; Wood, Neels, & Vergauwen, 2016), which typically masks considerable variation at the subnational level. In tandem with the growing availability of regional data, research on the childcare-fertility hypothesis progressively focuses on different regions within a single country (Hank & Kreyenfeld, 2003; Rindfuss et al., 2007), which limits the amount of (unobserved) confounding factors in comparison to cross-national studies as all country-specific characteristics are fixed (Thevenon, 2011). Second, several literature reviews identify the insufficient acknowledgement of individual uptake as a potential source of bias (Gauthier, 2007; Neyer & Andersson, 2008). Analyses of the impact of family policies on fertility should primarily focus on those who are likely to make use of the policy. Third, besides causal effects, associations between childcare availability and fertility may also occur due to spurious associations related to unobserved or omitted variables, reverse causation where fertility stimulates childcare provision, and selective migration which inflates fertility levels in contexts with extensive childcare services (Neyer & Andersson, 2008; Rindfuss et al., 2007; Thevenon, 2011). Although these issues have been well-documented, few empirical studies have aimed to remediate them.

As a result this study uses detailed longitudinal census and register data for the 2000s in Belgium combined with childcare coverage rates for children aged under three at the municipality-level, to assess the relation between local childcare availability on the one hand and first, second and third births among dual earner couples on the other. In addition, in order to narrow down the range of explanations for the association between childcare availability and fertility, we implement both random (multilevel) and fixed-effects models, and perform sensitivity analyses for reverse causation and selective migration. Finally, this study also quantifies the magnitude of the effect of local childcare availability on aggregate-level fertility.

In addition to data availability, Belgium is an interesting country to study the effect of childcare on fertility for two reasons. First, Belgium is characterized by a long history of reconciliation policies and

is, alongside France and Scandinavian countries, regarded as a context in which work and family are relatively compatible (Gornick, Meyers, & Ross, 1997; Klüsener, Neels, & Kreyenfeld, 2013; Matysiak & Węziak-Białowolska, 2016; Neyer, 2003). Since the 1970s formal childcare has been continuously extended and as of the early 2000s, Belgium is included in a short list of countries that meet the Barcelona target of 33 per cent childcare enrolment for children aged 0-3 (Farfan-Portet, Lorant, & Petrella, 2011; Morel, 2007; Plantenga et al., 2013; Population Council, 2006; Vandelannoote, Vanleenhove, Decoster, Ghysels, & Verbist, 2013). Second, a lack of effects of childcare coverage on fertility in previous studies for West-Germany and Sweden is respectively interpreted as a result of the overall inflexibility of the West-German childcare system (Hank & Kreyenfeld, 2003) and the general sufficiency of childcare provisions in Sweden (Andersson, Duvander, & Hank, 2003). Hence previous research suggest that research on regional childcare availability and fertility requires countries with meaningful variation in the local supply of childcare. Belgium is such a country as childcare availability and shortages vary considerably between regions, and the availability of local childcare places exhibits considerable increases during the early 2000s in line with explicit policy targets (Hedebouw & Peetermans, 2009; Kind & Gezin, 2000-2003; Plantenga et al., 2013; Vande Gaer, Gijselinckx, & Hedebouw, 2013; Vandelannoote et al., 2013).

## 2. The Belgian childcare context

Various different dimensions of formal childcare provision – price, quality, flexibility, availability – have been distinguished which potentially have different effects on fertility (Andersson, Duvander, & Hank, 2004). The cost of formal childcare has been identified as a major factor in countries with limited public childcare or limited price regulation like the US or the UK (Blau, 2001; Blau & Robins, 1989; Lehrer & Kawasaki, 1985; Mason & Kuhlthau, 1992), which contrasts with European countries with publicly organized and subsidized childcare (De Henau, Meulders, & O'Dorchai, 2007; Farfan-Portet et al., 2011; Gustafsson & Stafford, 1992). In Belgium subsidized and non-subsidized formal childcare coexist with about one quarter being non-subsidized (Farfan-Portet et al., 2011). The main direct subsidy depends on the amount of children and hours of care and average staff age. Additional subsidies are available for starting day care centres or centres with income-related tariffs, care for school age children or facilities for children with specific needs (Kind & Gezin, 2000-2003; ONE, 2018). Regional childcare institutions set the price for subsidized childcare, in contrast to free price-setting in the non-subsidized sector (Van Lancker & Ghysels, 2012). However, competition leads private childcare providers to adopt the maximum price among subsidized childcare (Farfan-Portet et al., 2011; Vandenbroeck, 2006). In addition to direct subsidies to childcare providers, Belgium also provides tax deductions for childcare expenses[2] (Farfan-Portet et al., 2011; ONE, 2018; Vandelannoote et al., 2013). A 2009 report for the Flemish region indicates that only 16.5 per cent of parents not using formal childcare claim that the cost is one of the reasons (Hedebouw & Peetermans, 2009).

Additionally, the quality of formal childcare provisions has been identified as an important dimension as people only use childcare if convinced that it is not harmful to their children (Andersson et al., 2004). All Belgian childcare providers need to be authorised by the regional childcare institutions imposing strict guidelines including dimensions of rooms and equipment, the child-staff ratio, staff level of education and future training, and medical tests for children (Dujardin, Fonder, & Lejeune, 2015; Farfan-Portet et al., 2011; Kind & Gezin, 2000-2003; ONE, 2018; Plantenga et al., 2013). Recent surveys among Belgian childcare users indicate that the overwhelming majority positively evaluates

[2] Since 2000 the maximum deductible sum is 11.2 Euros per day per child (Van Lancker and Ghysels 2012).

childcare quality (Hedebouw & Peetermans, 2009). Previous research also suggests that insufficient flexibility also hampers positive effects of formal childcare provision on fertility (Hank & Kreyenfeld, 2003). In Belgium, accessibility is one of the main requirements for subsidies. This implies that services must be open for at least 10 hours a day and 220 days a year (Dujardin et al., 2015). Non-users of formal childcare rarely mention reasons related to flexibility (Hedebouw & Peetermans, 2009).

This article argues that – compared to the cost, quality or flexibility - the local availability of formal childcare is more likely to affect fertility decisions. The presence of childcare centres depends strongly on regulations concerning subsidies, price-setting and quality requirements set by the federal and regional governments (Kind & Gezin, 2000-2003; ONE, 2018), and non-subsidized childcare exhibits high numbers of providers entering and exiting the market due to strong competition (Farfan-Portet et al., 2011). During the early 2000s, the period under consideration in this study, shortages have been documented, waiting lists occur, and research has reported on parents' failure to gain access to formal childcare (Farfan-Portet et al., 2011; Van Lancker & Ghysels, 2012; Vandelannoote et al., 2013). In contrast to free pre-primary education for children from 2.5 years on which is part of the educational system (Farfan-Portet et al., 2011; Van Lancker & Ghysels, 2012), there is no legal entitlement to childcare in Belgium as is the case in Sweden (Van Lancker & Ghysels, 2012). As a result of waiting lists and shortages in formal childcare at the turn of the century, regional and federal governments set a considerable increase in capacity as an explicit policy goal[3], increased subsidies, strengthened social security rights for childminders and introduced local childcare councils including all stakeholders to address shortages (Kind & Gezin, 2000-2003). Although childcare coverage exhibits a steady increase over the period considered in this study (2001-2004) (Dujardin et al., 2015; Farfan-Portet et al., 2011; Kremer, 2006), this increase does not surpass the increase in births nor the increasing female labour supply (Vandelannoote et al., 2013) and reports suggest that formal childcare cannot keep up with the growing demand (Hedebouw & Peetermans, 2009; Vande Gaer et al., 2013).

As local formal childcare is not always readily available, alternative strategies to combine work and family are also adopted, such as parental leave or informal childcare. During the early 2000s, Belgium exhibits a relatively flexible system of parental leave – which is an individual right depending on one's labour market position - with varying degrees of labour reduction: 100% for maximum three months, 50% for up to six months or 20% limited to 15 months (since 2002). Parents are allowed to split up the leave period depending on the sector of employment and previous work history and it can be used until the child is 4 years old[4] (Merla & Deven, 2013). Previous research indicates that, controlling for eligibility, only a minority of parents uses parental leave and the part-time options are by far the most popular (Anxo, Fagan, Smith, Letablier, & Perraudin, 2007; Kil, Wood, & Neels, 2018; Wood, Neels, Marynissen, & Kil, 2017). Finally, despite the fact that informal childcare is strongly decreasing in favour of formal care, a considerable share[5] of parents rely on grandparents as a primary source of childcare (Hedebouw & Peetermans, 2009). However, informal childcare for children aged under three is mostly used as a supplementary source of care in combination with formal childcare (Hedebouw & Peetermans, 2009; Vande Gaer et al., 2013).

[3] For instance, in 2000 the minister for social welfare in the Flemish government set the creation of 10 000 extra places in formal childcare as a policy goal (Kind & Gezin, 2000-2003).
[4] This age ceiling was extended to 6 years in 2005 and 12 years in 2009.
[5] In Flanders the share has decreased from 34.3 to 22.4 per cent in 2002-2009 (Hedebouw & Peetermans, 2009).

## 3. Theoretical and methodological considerations

### 3.1 The childcare-fertility hypothesis

The rise of the dual earner model stands central in both microeconomic and ideational theories of fertility decline in the developed world. According to Gary Becker's Nobel prize-winning New Home Economics (Becker, 1981), female labour force participation increases the opportunity costs of all non-market activities. Time-consuming activities such as childrearing entail costs in terms of forgone wages and the devaluation of skills, but also potentially jeopardize career tracks in the long run (Felmlee, 1995; Shapiro & Mott, 1994). In addition to the cost of time, literature also highlights potential socio-cultural role incompatibility between the roles of mother and worker. In contrast to pre-industrial societies, developed Western economies exhibit a clear separation between the workplace and the home, and women have been steered towards caring roles in the wake of industrialisation (Rindfuss & Brewster, 1996). The socio-cultural work-family tension for mothers stems from an increase in women's participation in employment in post-industrial societies which has not been accompanied by an equivalent shift away from social norms prescribing mothers as primary carers (Goldscheider, Bernhardt, & Lappegard, 2015; McDonald, 2000). For decades the work-family tension has been put forward as an explanation for fertility declines in Western countries. Although many contributions documented lower fertility for economically active women, there is no systematic evidence of lower fertility intentions for this group (Beaujouan, Sobotka, & Brzozowska, 2013; De Wachter & Neels, 2011; Ruokolainen & Notkola, 2002). This discrepancy was called the "fertility gap".

The assumption that the work-family tension depends on the time and place considered has long been supported in the literature (Becker, 1981; Liefbroer & Corijn, 1999; Rindfuss & Brewster, 1996). In addition to authors that put forward changing gender roles as the key to relieve the work-family tension for women (Goldscheider et al., 2015) or studies emphasizing the importance of intergenerational solidarity and informal childcare (Thomese & Liefbroer, 2013), literature exhibits a long-standing interest in state support through work-family reconciliation policies such as formal childcare (Blau & Robins, 1989; Rindfuss, Guilkey, Morgan, & Kravdal, 2010). The local availability of childcare is expected to positively affect dual earners' fertility. From an economic perspective, the possibility to outsource childrearing tasks lowers the opportunity costs of childbearing (Becker, 1981; Raz-Yurovich, 2014). From a socio-cultural perspective, it stimulates a context in which the role of the state as care provider is accepted and the combination of work and family is approved (Fagnani, 2002; Gauthier, 2007; Mills, Rindfuss, McDonald, & Te Velde, 2011; Sjöberg, 2004). Given the aforementioned economic and socio-cultural impact of local formal childcare provision on the work-family tension, we expect dual earner couples to be more likely to have a child in case formal childcare is readily available (hypothesis 1).

### 3.2 specifying childcare effects

Available literature documents many theoretical and methodological challenges regarding the specification of the effect of local childcare provision on fertility.

#### *3.2.1 Population heterogeneity*

Literature reviews have identified the insufficient acknowledgement of population heterogeneity in the uptake and effects of work-family reconciliation policies as a potential explanation for the inconclusive findings regarding the effect on childbearing (Gauthier, 2007; Neyer & Andersson, 2008).

Assessments should primarily focus on individuals or households experiencing work-family tension and who's fertility is expected to respond to local variation in childcare availability. As a result, this study focuses on dual earner couples. Although it should be acknowledged that local childcare availability may also impact fertility of other household types directly or indirectly through labour market positions, it is more cumbersome to hypothesise about these relations as non-employed parents can take up caregiving and the magnitude and direction of the associations between labour market positions and fertility have been found to vary depending on the population subgroup considered (Wood & Neels, 2017).

*3.2.2 Parity*

Available research is inconclusive on whether a different impact of childcare availability can be expected depending on the birth order considered. Conflicting hypotheses can be thought of, depending on how the context-contingency of fertility is conceptualized. In the first view, the timing of first births is most dependent on the childcare context, whereas the timing of higher-order births is less flexible as dual earner couples want to provide the firstborn a playmate, approach the biological age-limits to fertility, or want to limit the time spent out of the labour force (Ni Bhrolchain & Beaujouan, 2012; Wood & Neels, 2017). To the extent that the occurrence of a first birth for dual earners signals that a manageable work-family strategy has been found, higher-order fertility is less differentiated by the local availability of formal childcare (hypothesis 2a). However, in the occurrence of low formal childcare availability, it is easier to find an alternative strategy for the care of one child (e.g. informal care) compared to several children and childless individuals have very little direct experience with work-family tension, waiting lists and supply shortages. As a result the alternative hypothesis is that particularly higher order births will be influenced by local childcare availability (hypothesis 2b) (Kravdal, 1996).

*3.2.3 Causal versus alternative interpretations*

The degree to which reported associations between childcare availability and fertility result from spurious associations, reverse causation, or selective migration remains a major question in the literature (Neyer & Andersson, 2008; Thevenon, 2011). With respect to spuriousness, the inability to control for other local characteristics which determine both childcare provision and fertility is identified as a source of bias (Baizan, 2009; Rindfuss et al., 2007). For instance, few labour market opportunities for young women deflate the demand for childcare, but also may lower fertility among dual earner couples. As a result, analyses relating childcare availability to fertility should include adequate controls for local characteristics. An additional source of spuriousness may arise when solely focussing on dual earner couples. It is likely that the local availability of formal childcare has a positive impact on the amount of dual earner couples. To the extent that a larger group of dual earner couples in response to higher childcare availability is selective in terms of childbearing, dual earner fertility will be affected due to self-selection. Hence research designs are required in which the sample of dual earner couples is not allowed to vary as a function of childcare availability. This study creates such a research design by selecting dual earner couples at a time point preceding the prospective follow-up of fertility patterns.

With respect to reverse causation, regions with high fertility may consequently also exhibit higher levels of childcare supply as a result of demand-driven childcare allocation (Castles, 2003; Mills et al., 2011; Thevenon, 2011). This generates a feedback effect, or so-called reverse causality, which may be

mistaken for a causal effect of childcare on dual earners' fertility. However, it should be noted that higher fertility by definition deflates coverage rates if the amount of places remains stable. Hence feedback effects only occur when increases in childcare places in response to higher fertility exceed the increase in fertility.

Finally, selective migration may also impact the relation between local childcare availability and dual earners' fertility (Rindfuss et al., 2007). Given that dual earner couples experience a strong work-family tension in the absence of childcare, municipalities with extensive local childcare provision are likely to attract dual earner couples with higher childbearing intentions. Although this potential mechanism has not been documented for Belgium, the assessment of the childcare-fertility nexus requires controls for internal migration, as selective migration can inflate dual earner fertility in the region of destination.

#### 3.2.4 The aggregate-level relevance of childcare effects

Although empirical analyses of how work-family policies may impact fertility decisions are routinely justified by referring to the context of low or even lowest-low fertility in the developed world, studies on the impact of local childcare availability on childbearing rarely attempt to quantify the impact of childcare availability on fertility at the aggregate level (Rindfuss et al., 2010). This study assesses the individual- and aggregate-level impact of local childcare availability on fertility. Using the estimates of models at the individual level, we calculate the effect of local childcare availability on fertility at the aggregate-level, taking into account direct effects of childcare availability on birth hazards, but also indirect effects via lower order births which determine the risk set for higher order births.

## 4. Data and Methods

### 4.1 Data

This study draws on 2001 Belgian census data linked to 2002-2005 data from the National Register. The 2001 Belgian census provides detailed information on all individuals legally residing in the country - including fertility histories, labour market position, education, nationality and marital status – and allows to identify heterosexual co-residential dual-earner couples (Deboosere & Willaert, 2004). The prospective research design of the linked census and register data allows us to study fertility during the period 2002-2005 among dual earner couples in 2001. Our sample consists of 157,476 couples at risk of a first birth, 216,331 couples at risk of a second birth, and 321,576 couples at risk of a third birth. These couples are followed from October 2001 until they experience a birth, or are censored when the female partner reaches the age of 50, when one of both partners dies or emigrates, or when the observation ends in January 2006. This results in a sample of 5,096,609 couple-months for the transition to parenthood, 6,992,385 couple-months for second births and 11,191,123 couple-months for third births. This prospective individual-level dataset also includes longitudinal information on the municipality of residence, which allows to combine the microdata with municipality-level data on the amount of places in formal childcare, provided by regional childcare institutions.

### 4.2 Random- and fixed-effects hazard models

Discrete-time hazard models for first, second and third births are estimated using a logit link function. As a result, exponentiated parameter estimates can be interpreted as odds-ratios. The hazard models were estimated separately by birth order for two reasons. First, this allows us to minimize the amount

of interactions while enabling order-specific effects for all covariates. Second, although the joint modelling of subsequent births in a shared-frailty model has been identified as a useful approach to better capture time-constant unobserved heterogeneity (Wood et al., 2014; Wooldridge, 2002), this approach is not feasible in this study since we observe maximum one birth per couple (Rindfuss et al., 2007).

As a result of the hierarchical nature of combined municipality- and couple-level data, simple hazard models of births at the micro level may generate biased estimates of the effect of local childcare availability on birth hazards. Previous research indicates two strands of research in this respect: random effects, or so-called 'multilevel' models, and municipality fixed effects models (Andersson et al., 2004; Hank & Kreyenfeld, 2003; Rindfuss et al., 2007). These two types of models exhibit clear differences in terms of which variation in childcare coverage is exploited, but also with respect to controlling for spuriousness in the relation between local childcare availability and dual earner fertility.

First, both types of models exploit different types of variation. Random effects models draw on variation between regional entities as well as over time. As a result, the occurrence of a positive association between local childcare availability and dual earner fertility in such a model implies that contexts – albeit different places or different times – with higher childcare availability also exhibit higher dual earner fertility. As we consider childcare coverage in a relatively short time period (2001-2004) variation in childcare availability between regions is much larger than changes over time. Hence the random effects model mostly exploit variation across municipalities. This contrasts with fixed effects models in which all differences between municipalities are controlled for by including municipality-dummies in the model. As a result, a positive association between local childcare availability and dual earner fertility implies that, within a municipality, times in which childcare coverage is higher will also exhibit higher dual earner fertility. It is likely that the experience of change in the local availability of childcare is a strong determinant of fertility decisions as one has experienced the situation both before and after the change. Similar arguments have been used to focus on changes in other contextual factors such as unemployment (Sobotka, Skirbekk, & Philipov, 2011).

Second, random and fixed effects models differ in the degree to which the endogeneity of local childcare availability is controlled for. Although random effects models yield unbiased parameter standard errors, these models do not control for municipal characteristics, unless included as variables into the model (Allison, 2009). The fixed-effects approach, which amounts to the inclusion of $(n-1)$ dummies for the regional units, does not require the identification of all relevant third variables which affect regional variation in childcare availability and fertility as only variation within municipalities over time is considered. Fixed-effects models use every municipality as its own control, making the identification of time-constant municipal characteristics which may render childcare availability endogenous unnecessary. However, even in a fixed-effects approach the association between childcare provision and fertility is not free of spuriousness as variability over time in childcare and fertility may still result from time-varying confounding factors. As a result, time-varying municipality-level covariates can be used in the model (Allison, 2009; Rindfuss et al., 2007).

Third, random and fixed effects models also differ in the extent to which the association between local childcare coverage and dual earner fertility may be driven by a positive effect of childcare coverage on the number of dual earners. A positive association between childcare coverage and dual earner fertility across municipalities – as exploited in the random effects model – may occur due to larger and possibly more positively selective groups of dual earner couples in terms of fertility, under higher childcare availability. This mechanism - in which the availability of childcare attracts couples with higher fertility intentions into the dual earner model, rather than affecting fertility among dual

earner couples – can also take place over time within a municipality. However, due to the fact that our prospective research design selects dual earner couples in 2001, regardless of their labour market positions in the 2002-2005 follow-up period, the population of dual earner couples is not allowed to vary as a function of childcare availability. Hence in the fixed-effects models the association between local childcare availability and dual earner fertility within municipalities over time cannot be attributed to the impact of childcare availability on the dual earner model. The other side of the same coin is that labour market positions are not observed longitudinally and to some extent the reported associations between childcare availability and fertility may be due to indirect effects via changing labour market positions. However, given that a decomposition of direct effects of childcare coverage on fertility and indirect effects via other variables lies beyond the scope of this article, this is not considered a limitation.

Benefitting from the availability of longitudinal data on local childcare coverage, we present both random-effects (Model A) and fixed-effects models (Model B). Several pairs of random-effects and fixed-effects models are estimated. Model 1 studies the effect of local childcare availability on birth hazards. Model 2 studies how local childcare availability shapes the height and form of fertility schedules in order to provide valid estimates of the impact on fertility at the aggregate level. Using varying effects by exposure, observed birth hazard functions are compared to the hazard functions which have been subjected to average marginal effects. Following previous research (Neels, 2006; Wood et al., 2014), this article presents synthetic parity progression ratios (SPPR) based on the birth hazard functions (equation 1).

$$SPPR_i = 1 - \prod_{t=0}^{30} (1 - \hat{q}_{(i,t)}) \tag{1}$$

This implies that the SPPR for birth order i is calculated using the estimated birth hazards ( $\hat{q}$ ) for the first 30 years of exposure (t). In addition, an SPPR-based indicator for combined first, second and third births is calculated (equation 2) to assess the impact of local childcare coverage on total fertility.

$$SPPR - based - TFR_{1-3} = SPPR_1 + (SPPR_1 * SPPR_2) + (SPPR_1 * SPPR_2 * SPPR_3) \tag{2}$$

4.3 Variables

The main independent variable of interest is local childcare coverage which equals the amount of places divided by the population aged 0-3. This variable is included with a 12 month time lag[6] to approximate the time of birth decisions. Since variation in childcare coverage over time within municipalities is limited compared to variation between municipalities, the random-effects models consider a 10 per cent difference in childcare coverage across place and time, whereas the fixed-effects models estimate the effect of a 1 per cent change over time.

At the couple level, we control for the following covariates: i) exposure, ii) the female partner's age at entry into the risk set, iii) education, iv) work regime, v) age of the male partner, vi) marital status, vii) origin group and viii) the availability of informal childcare. *Exposure* (time-varying) equals months

[6] As it is possible that the lag between childcare availability and fertility decisions is larger, additional analyses (not shown) have been performed using 24 or 36 month time lags. These do not change the main results.

since her graduation for first births and months since her previous birth for higher-order births. Months since graduation is used as exposure variable instead of months since her fifteenth birthday, as very few women in the youngest age groups are in a dual earner couple, entailing unstable hazard functions at early exposures (Wood & Neels, 2017). For first births a cubic polynomial baseline function is used For second births a spline function with one node at 36 months is used and third birth hazard functions are best approximated by a spline function with nodes at 24, 36 and 48 months. To control for *age at entry into the risk set* (time-constant), we control for age at graduation, age at first birth and age at second birth in models for first, second and third births respectively, including linear and quadratic terms. *Education* in 2001 (time-constant) distinguishes low education (ISCED 0-2) from medium education (ISCED 3-4) and high education (ISCED 5-6) for both partners. All nine couple combinations are included in the models. We also include interactions between all exposure variables and couples' educational attainment. *Work regime* in 2001 (time-constant) distinguishes full-time employment from part-time employment. All four couple combinations are included in the models and the baseline hazard function is interacted with couples' work regime. *Age of the male partner* (time-varying) is included as a categorical variable with six categories: younger than 25, 25-29, 30-34 (reference category), 35-39, 40-45, and older than 45. *Marital status* in 2001 (time-constant) distinguishes married couples from non-married couples. *Origin group* (time-constant) is based on respondents' and their parents' nationality at birth and distinguishes seven groups: i) Belgian, ii) neighbouring country (United Kingdom, France, the Netherlands, Luxemburg, Germany), iii) Southern European countries, iv) other European countries, v) Turkey or Morocco, vi) non-European highly developed countries (The United States, Canada, Japan, Australia, New Zealand), and vii) other non-European countries. Both her and his origin group are included in the models. To proxy the *availability of informal childcare*, we include a time-constant dummy indicating whether there is a non-employed adult in the household in 2001. In addition we include time-varying information on whether partners were born in the municipality of residence. All four couple combinations are included in the model and a couple of non-locally born individuals is the reference category.

In contrast to the fixed-effects models, the random-effects models include additional municipality-level variables which may capture the endogenous placement of childcare facilities. In order to capture rural-urban differences, we control for *population density* distinguishing between less than 250, 250-499, 500-749, 750-1500, and more than 1500 inhabitants per square kilometre. *Female opportunities* in the labour force could affect both childcare provisions and regional fertility patterns (Baizan, 2009; Rindfuss et al., 2007). However, previous studies - including women's labour force participation into regression models - note that these variables may reflect maternal employment and childcare availability (Baizan, 2009). Consequently this article uses female labour force participation for childless women between 18 and 49 in 2001 derived from the Belgium 2001 census. Due to space restrictions, the estimates for the control variables are provided in an online appendix (table A1).

### 4.4 Robustness checks

Finally, a series of sensitivity analyses tests the robustness of our results for selective migration, reverse causality, couple dissolution, non-linear childcare effects, alternative clocks for first births, and full-time versus part-time employment. With respect to selective migration, we estimate models using only couples who have been living in the same municipality for at least three years. In order to control for feedback mechanisms between fertility and childcare availability, we include local total fertility

rates or birth numbers lagged by 5-10 years (Rindfuss et al., 2007). In addition, given that couple dissolution severely restrains childbearing hazards, we estimate models excluding couples that separated by 2006. To assess non-linear effects of childcare availability on dual earner fertility, models including quadratic or categorical effects are estimated. As an alternative to duration since graduation, duration since age 15 is also used as a clock for first births. Finally, varying effects for couples in which one or both partners works part-time are assessed by including interactions.

## 5. Results and Discussion

### 5.1 Local childcare availability in Belgium

During the time period considered in this study, 2001-2004, Belgium exhibited considerable variation in the availability of formal childcare between municipalities, as well as variation within municipalities over time. Figure 1 displays regional variation in childcare coverage and indicates that in general the Northern part of the country, Flanders, exhibits a higher availability of formal childcare. During the mid-2000s, Belgium exhibits a forerunner position being one of six countries meeting the Barcelona target of childcare for children under three, and especially the Flemish region was at the forefront of the childcare expansion (Population Council, 2006). Additionally, it should be noted that besides this north-south divide, all regions – Flanders, Wallonia and Brussels - display strong variability in formal childcare coverage between municipalities.

Figure 2 visualizes within municipality variation in childcare coverage over time. It is clear that this variation is more limited. Most municipalities witness an increase of the coverage rate and the average evolution from 2001 to 2004 is a 2.95 per cent rise, which corresponds to explicit policy goals to increase the amount of places in times of high demand and rising fertility (Kind & Gezin, 2000-2003; Vandelannoote et al., 2013). As positive and negative evolutions cancel each other in the calculation of the average, the average of the evolutions in absolute values is also calculated, indicating an average change of 4.19 per cent in the 2001-2004 time period. In the context of policy-makers attempts to increase childcare availability and combat childcare shortages and long waiting lists, this magnitude of change potentially affects fertility decisions.

*FIGURES 1 & 2 ABOUT HERE*

### 5.2 Local childcare availability and fertility

#### *5.2.1 Observed birth hazard schedules by childcare coverage*

A descriptive yet first indication of the degree to which birth hazards for dual earner couples are associated with childcare coverage at the municipality level is provided in Figure 3. Comparing the birth schedules by childcare coverage suggests that birth hazards are not indifferent to the local availability of formal childcare. The parts of the fertility schedules characterized by the highest birth hazards suggest unambiguous positive associations with childcare coverage for all birth orders. This pattern seems to be relatively linear for all births, the low third birth hazards for contexts with a childcare coverage rate of 40-50 per cent being the exception. However, first birth schedules also indicate weak negative relations between childcare coverage and birth hazards among couples in which the female

partner graduated either recently or a relatively long time ago. Although these descriptive results do not provide any explanations for the observed patterns, a potential explanation for the negative bivariate association between childcare availability and first birth hazards close to graduation is that contexts with high childcare availability allow for women to invest strongly in their career before relying on formal childcare, entailing later transitions to parenthood (Liefbroer & Corijn, 1999). For second and third births a clear positive association is found between birth hazards and local childcare coverage in the first three to five years since the previous birth. In contrast, later sections of the schedules often indicate a negative association between childcare coverage and birth hazards. A potential explanation for this pattern is that the local availability of formal childcare speeds up higher order childbearing. However a true distinction between tempo and quantum effects lies beyond the scope of this 2002-2005 period analysis.

*FIGURE 3 ABOUT HERE*

*5.2.2 Associations over place and time*

In order to test our expectation that dual earner couples will be more likely to have a child in case formal childcare is readily available, we rely on multivariate associations between local childcare coverage and birth hazards (Table 1). Model 1A, a random-effects model, indicates a significantly positive association between local childcare coverage and fertility over place and time, which is most pronounced for first births. A 10 per cent difference in formal childcare coverage is related to a ((1.161-1)*100) 16.1 per cent difference in the odds of having a first birth, whereas the corresponding increase for higher-order births is limited to 5.1 per cent. A larger, though frequently observed, difference of 25 per cent in the coverage rate of formal childcare is related to a ($(1.161^{2.5}-1)*100$) 45.2 per cent difference in the odds of having a first birth and a 13.2 per cent difference in the odds of having a higher-order birth. Although these random-effects estimates are not contaminated due to the inclusion of groups which are unlikely to use formal childcare (e.g. inactive couples) and the impact of omitted variable bias is minimised by the inclusion of a wide range of control variables (Gauthier, 2007; Neyer & Andersson, 2008), it is likely that the reported associations are biased by the endogenous placement of formal childcare places and the impact of local childcare availability on the dual earner model (Baizan, 2009; Rindfuss et al., 2007).

As a result, Model 1B exploits variation in childcare coverage within municipalities over time using fixed-effects regressions which have been identified as a superior approach to limit spurious associations (Allison, 2009; Baizan, 2009; Rindfuss et al., 2007). Using every municipality as its own control, changes in childcare coverage within a municipality over time are significantly and positively associated with birth hazards. In line with the results from random-effects models the strongest effect is found for first births. A one per cent increase in formal childcare coverage in a given municipality is related to a ((1.108-1)*100) 10.8 per cent increase in first birth odds whereas the corresponding increase for higher-order births is limited to 2.8 per cent and 2.1 per cent for second and third births respectively. Three key differences between the fixed- and random-effects model should be taken into account when comparing the effects of both models. First, fixed-effects models assess the degree to which change in local childcare availability associates with dual earner fertility. It is likely that the experience of improving childcare provision in a given context has a strong impact on couples' fertility decisions. Second, fixed-effects model estimates are not biased by the potentially endogenous placement of local childcare over Belgian municipalities. Third, the random-effects model estimates are potentially biased by the impact of local childcare availability on the size and selectivity of the

group of dual earner couples in a given municipality. This is not the case in the fixed-effects model as our prospective research design selects dual earner couples in 2001, regardless of their labour market positions in the 2002-2005 follow-up period. Hence 2002-2005 fertility changes within a municipality cannot be due to the increased size and selectivity of the group of dual earners.

Finally, six additional robustness checks were performed, none of which notably altered the associations between childcare coverage and birth hazards[7]. First, local childcare availability could also be endogenous due to reverse causation to the extent that past fertility trends affect childcare provision (Thevenon, 2011). However the inclusion of local number of births or fertility rates with time lags ranging from five to ten years yield similar results (Rindfuss et al., 2007). Second, municipalities with extensive local childcare provisions are likely to attract dual earner couples who intend to have a birth, in turn inflating dual earner fertility in the destination municipality. However, very similar associations between local childcare coverage and birth hazards are found when restricting the sample to dual earner couples exhibiting residential stability for at least three years. Third, when limiting the sample to couples which were still living in the same household in 2006, the main findings were not altered. Fourth, similar results are found when months since the fifteenth birthday is used as an alternative clock for first births. Fifth, the inclusion of interactions between work regime (full-time or part-time) and childcare availability did not yield any significant improvement of the model. Finally, the assessment of non-linear effects indicated weak non-linear curves. In addition, categorical effects of local childcare availability show that the only clear deviation from a linear pattern is the lower third birth hazards in the second highest childcare availability category, which is in line with the descriptive findings in figure 3.

*5.2.3 The impact on aggregate-level fertility*

In order to provide accurate estimates of how the local availability of formal childcare affects fertility at the aggregate level, we estimate models allowing an interaction between exposure and childcare coverage (Model 2A-B). Since the interpretation of differential effects over the birth schedule in terms of odds ratios is cumbersome, marginal effects of childcare coverage on birth hazards by year of exposure are provided in appendix (figure A1). The random-effects models indicate that the inclusion of interactions between exposure and childcare coverage yield a significant model improvement for the transition into parenthood ($\Delta$-2LL= 75.5, $\Delta$df= 3, p< .001), second ($\Delta$-2LL=286.1, $\Delta$df=2, p<.001) and third births ($\Delta$-2LL=110.622, $\Delta$df= 4, p<.001). Despite limited significant negative effects at long durations, all effects of childcare coverage on first birth hazards are significantly positive. For higher-order births a different picture emerges as, besides positive effects early after the previous birth, noteworthy significant negative effects occur at long durations. Similarly, fixed-effects models including the interaction between exposure and childcare coverage show significant improvements for first ($\Delta$-2LL=78, $\Delta$df=3, p<.00), second ($\Delta$-2LL=329.4, $\Delta$df=2, p<.00), and third births ($\Delta$-2LL=117.5, $\Delta$df=4, p<.00). The most noteworthy difference compared to the random effects models is that the marginal effects of childcare coverage in the fixed-effects models show no negative effects for first and third births, and only very limited negative effects on second birth hazards.

To quantify the impact of local childcare availability on aggregate-level fertility, we compare fertility indicators (Table 2) based on the observed birth schedules to indicators calculated based on schedules subjected to the marginal effects in Models 2A-B (see figure A2, online appendix). The positive impact of local childcare availability on the SPPR is substantial for first births. The observed first birth function

[7] Results are not presented here, but available upon request.

for Belgium yields a $SPPR_1$ of 0.895, whereas the application of the random-effects estimates of the impact connected to a 10 per cent rise in childcare coverage entails a $SPPR_1$ of 0.921 and the fixed effects impact of a 1 per cent increase in childcare coverage over time implies a $SPPR_1$ of 0.915. For higher order births, the impact in terms of SPPR is more limited. Compared to the 0.680 observed $SPPR_2$, a 10 per cent higher childcare coverage rate is related to a $SPPR_2$ of 0.692 and a 1 per cent increase in childcare coverage within municipalities over time yields a $SPPR_2$ of 0.689. Similarly the observed $SPPR_3$ is 0.288, whereas a 10 per cent higher childcare coverage rate is related to a $SPPR_3$ of 0.301 and a 1 per cent increase in childcare coverage within municipalities over time yields a $SPPR_3$ of 0.293.

However, in addition to the direct effect of local childcare coverage on birth hazards, the number of higher order births also depends on the amount of couples who enter the risk set by having the previous birth. The observed SPPR-based-$TFR_{1-3}$ for Belgium as a whole indicates a total of 1.679 children per dual earner couple, whereas the corresponding value under a 10 per cent change in childcare coverage from the random-effects model is 1.750 children per women, and the value for a 1 per cent increase in the fixed-effects model is 1.730 children per women. These 4.2 and 3.0 per cent changes in the total number of first, second and third births combined seem considerable given the low and even lowest fertility levels in the developed world.

*TABLE 1-2 ABOUT HERE*

## 6. Conclusion

The theoretically well-grounded hypothesis that local formal childcare availability positively affects childbearing has been part of the literature for a long time. The increasing availability of formal childcare in many developed countries is assumed to reconcile work and family life as high opportunity costs of childrearing tasks are reduced and mother and worker roles become more compatible (Brewster & Rindfuss, 2000; Rindfuss & Brewster, 1996; Rindfuss et al., 2007). Despite these theoretical underpinnings, researchers differ in their opinion on whether family policy impacts fertility behaviour (Demeny, 2003; Rindfuss et al., 2007), and previous empirical studies on the association between childcare availability and fertility have provided ambiguous results (Andersson et al., 2004; Baizan, 2009; Del Boca, 2002; Gauthier, 2007; Hank & Kreyenfeld, 2003; Neyer & Andersson, 2008; Rindfuss et al., 2010; Rindfuss et al., 2007; Wood et al., 2016).

This article argues that these mixed findings occur - at least partly - due to variation in the methods applied (Gauthier, 2007; Neyer & Andersson, 2008). We contribute to the literature by exploiting rich longitudinal Census and Register data for 588 municipalities in Belgium, which has consistently been at the forefront of the formal childcare expansion (Klüsener et al., 2013; Population Council, 2006). Belgium also exhibits considerable variation in formal childcare coverage over regions and time during the early 2000s, a period in which the expansion of formal childcare and combatting long waiting lists were explicit policy targets (Farfan-Portet et al., 2011; Kind & Gezin, 2000-2003).

This study adds to the growing literature on the effect of formal childcare availability on fertility using subnational variation in childcare coverage (Andersson et al., 2004; Baizan, 2009; Del Boca, 2002; Hank & Kreyenfeld, 2003; Rindfuss et al., 2007). Following Neyer and Andersson (2008) who state that policy evaluations should limit their sample to those who are affected by the policy, this study

considers dual earner couples to assess the effect of local formal childcare availability on fertility. Our findings indicate clear and substantial positive effects (hypothesis 1).

The positive effects on first and higher-order births reported here are consistent with previous findings for Norway (Rindfuss et al., 2010; Rindfuss et al., 2007), Spain (Baizan, 2009), and Italy (Del Boca, 2002), but contradict findings for West-Germany (Hank & Kreyenfeld, 2003) and Sweden (Andersson et al., 2004). Although varying findings between countries may occur as a result of country-specific factors, such as the overall inflexibility of West-German formal childcare (Hank & Kreyenfeld, 2003) and the general sufficiency of childcare provisions in Sweden (Andersson et al., 2003), it is noteworthy that previous studies variably use multilevel models or fixed-effects regression techniques. In order to document the sensitivity of the association between local childcare coverage and birth hazards, this paper benefits from the availability of longitudinal information on local childcare coverage to compare random-effects and fixed-effects hazard models. Significant positive associations in the random-effects model imply that variation in childcare coverage over place and time is positively related to fertility. However, the inclusion of a random municipality effect does not control for unobserved heterogeneity at the municipality level (Allison, 2009) and has been criticized as a method to study the impact of local childcare availability on fertility (Rindfuss et al., 2007). As a result, this study also presents fixed-effects models, in which every municipality is used as its own control variable and only variation in childcare coverage within a municipality over time is considered. These models indicate that variation in childcare coverage and childbearing within municipalities across time are also positively associated.

However, when allowing the effect of local childcare coverage on birth hazards to vary by exposure, only random-effects models indicate negative effects of childcare coverage at later exposures. This finding is in line with results for Norway indicating that a negative effect of childcare on first births reverses to a positive effect when including region fixed effects (Rindfuss et al., 2007). Under the assumption that the causal effect of childcare availability on dual earner fertility is positive, this difference in results suggests that associations resulting from fixed-effects models are less sensitive to spuriousness (Allison, 2009; Baizan, 2009; Rindfuss et al., 2007). In addition to the use of fixed-effects models, additional sensitivity analyses indicate that the positive association between local childcare coverage and birth hazards is robust to reverse causation and selective migration (Neyer & Andersson, 2008; Rindfuss et al., 2007; Thevenon, 2011).

Both random- and fixed-effects models indicate that the positive association between local childcare availability and dual earner fertility is strongest for first births. This suggests that the timing of parenthood is most contingent on contextual factors such as childcare availability, whereas the timing of higher-order births is less flexible as dual earner couples want to provide the firstborn a playmate, approach the biological age-limits to fertility, or want to limit the time spent out of the labour force (hypothesis 2a) (Ni Bhrolchain & Beaujouan, 2012; Wood & Neels, 2017).

In contrast to most available literature (see Rindfuss et al. 2010 for an exception), this study also assesses the impact of childcare coverage on aggregate-level fertility. Our results indicate that parity progression ratios and the number of combined first, second and third births per women increase substantially in response to higher childcare coverage rates over place and time. Hence these findings support the view that formal childcare services can be a fruitful tool in the context of low and lowest-low fertility levels in the developed world.

Finally, six limitations and avenues for future research are identified. First, the prospective register data used in this study does not provide longitudinal information on labour market positions. Hence the reported associations between childcare and fertility may partly reflect indirect effects via changing labour market positions. Second, notwithstanding our focus on dual earners and thus population heterogeneity, much work remains to be done to analyse formal childcare effects for other types of households (Puur et al., 2016; Van Bavel & Rozanska-Putek, 2010; Wood et al., 2016). Third, as a result of the inability to locate dual earners' parents and other family members, this study relies on less preferable indicators for the availability of informal childcare. Fourth, although anecdotal evidence suggests that considerable numbers of parents use formal childcare in the region of employment, the documentation of such strategies and effects on fertility lies beyond the scope of this study. Fifth, although sensitivity models for couples that are still cohabiting in 2006 are performed, a lack of union histories hampers censoring at the time of separation. Finally, a distinction between timing and quantum of fertility lies beyond the scope of this period analysis. This requires multiple decades of observations of local childcare availability in tandem with individual fertility data, allowing both period and cohort fertility to be assessed.

## 7. References


Allison, P. (2009). *Fixed Effects Regression Models (Quantitative Applications in the Social Sciences)*: SAGE Publications.

Andersson, G., Duvander, A.-Z., & Hank, K. (2003). Do Child Care Characteristics Influence Continued Childbearing in Sweden? An Investigation of the Quantity, Quality, and Price Dimension. *MPDIR Working paper WP 2003-013*.

Andersson, G., Duvander, A.-Z., & Hank, K. (2004). Do child-care characteristics influence continued child bearing in Sweden? An investigation of the quantity, quality, and price dimension. *Journal of European Social Policy, 14*(4), 407–418.

Anxo, D., Fagan, C., Smith, M., Letablier, M.-T., & Perraudin, C. (2007). Parental Leave in European Companies. Dublin: European Foundation for the Improvement of Living and Working Conditions.

Baizan, P. (2009). Regional child care availability and fertility decisions in Spain. *Demographic Research, 21*(27), 803-842.

Beaujouan, E., Sobotka, T., & Brzozowska, Z. (2013). *Education and sex differences in intended family size in Europe, 1990s and 2000s*. Paper presented at the Changing families and fertility choices, Oslo, Norway.

Becker, G. (1981). *A Treatise on the Family*. London: Harvard University Press.

Blau, D. M. (2001). *The Child Care Problem: an Econometric Analysis*. New York: Russel Sage.

Blau, D. M., & Robins, P. K. (1989). Fertility, employment, and childcare costs. *Demography, 26*(2), 287-299.

Brewster, K. L., & Rindfuss, R. R. (2000). Fertility and women's employment in industrialized nations. *Annual Review of Sociology, 26*, 271-296.

Castles, F. G. (2003). The world turned upside down: below replacement fertility, changing preferences and family-friendly public policy in 21 OECD countries. *Journal of European Social Policy, 13*(3), 209-227.

De Henau, J., Meulders, D., & O'Dorchai, S. (2007). Making time for working parents: Comparing public childcare provision. In D. Del Boca & C. Wetzels (Eds.), *Social Policies, Labour Markets and Motherhood. A Comparative Analysis of European Countries* (pp. 28-62). Cambridge: Cambridge University Press.

De Wachter, D., & Neels, K. (2011). Educational differentials in fertility intentions and outcomes: family formation in Flanders in the early 1990s. *Vienna Yearbook of Population Research, 9*, 227-258.

Deboosere, P., & Willaert, D. (2004). *Codeboek algemene socio-economische enquête 2001*. Working Paper 2004-1. Steunpunt Demografie Vakgroep Sociaal Onderzoek (soco) Vrije Universiteit Brussel. Brussels.

Del Boca, D. (2002). The effect of child care and part time opportunities on participation and fertility decisions in Italy. *Journal of Population Economics, 15*(3), 549-573.

Demeny, P. (2003). Population Policy Dilemmas in Europe at the Dawn of the Twenty-first Century. *Population and Development Review, 29*, 1-28.

Dujardin, C., Fonder, M., & Lejeune, B. (2015). Does formal child care availability for 0-3 year olds boost mothers' employment rate? Pandel data based evidence from Belgium. *IWEPS working paper*.

Fagnani, J. (2002). Why do French women have more children than German women? Family policies and attitudes towards child care outside the home. *Community, Work & Family, 5*(1), 103-119.

Farfan-Portet, M.-I., Lorant, V., & Petrella, F. (2011). Access to childcare Services: The Role of Demand and Supply-side Policies. *Population Research and Policy Review, 30*, 165-183.

Felmlee, D. H. (1995). Causes and Consequences of Women's Employment Discontinuity 1967-1973. *Work and Occupations, 22*, 167-188.

Gauthier, A. H. (2007). The impact of family policies on fertility in industrialized countries: a review of the literature. *Population Research and Policy Review, 26*(3), 323-346.
Goldscheider, F., Bernhardt, E., & Lappegard, T. (2015). The Gender Revolution: A Framework For Understanding Changing Family and Demographic Behaviour. *Population and Development Review, 41*(2), 207-239.
Gornick, J., Meyers, M., & Ross, K. (1997). Supporting the employment of mothers: Policy variation across fourteen welfare states. *Journal of European Social Policy, 7*(1), 45-70.
Gustafsson, S., & Stafford, F. P. (1992). Child care subsidies and labour supply in Sweden. *Journal of Human Resources, 27*(1), 204-230.
Hank, K., & Kreyenfeld, M. (2003). A Multilevel Analysis of Child Care and Women's Fertility Decisions in Western Germany. *Journal of Marriage and Family, 65*(3), 584-596.
Hedebouw, G., & Peetermans, A. (2009). Onderzoek naar het gebruik van opvang voor kinderen jonger dan 3 jaar in het Vlaamse Gewest in 2009: HIVA – K.U.Leuven, Steunpunt Welzijn, Volksgezondheid en Gezin, Kind en Gezin.
Kil, T., Wood, J., & Neels, K. (2018). Parental leave uptake among migrant and native mothers: Can precarious employment trajectories account for the difference? *Ethnicities, 18*(1). doi: doi:10.1177/1468796817715292
Kind & Gezin. (2000-2003). Jaarverslagen Kinderopvang 2000-2003.
Klüsener, S., Neels, K., & Kreyenfeld, M. (2013). Family policies and the Western European Fertility Divide: Insights from a Natural Experiment in Belgium. *Population and Development Review, 39*(4), 587-610.
Kravdal, O. (1996). How the local supply of day-care centers influences fertility in Norway: A parity-specific approach. *Population Research and Policy Review, 15*, 201-218.
Kremer, M. (2006). The politics of ideals of care: Danish and Flemish child care policy compared. *Social Politics, 13*(2), 261-285.
Kreyenfeld, M., & Hank, K. (2000). Does the availability of child care influence the employment of mothers? Findings from western Germany. *Population Research and Policy Review, 19*(4), 317-337. doi: 10.1023/a:1026556309080
Lehrer, E. L., & Kawasaki, S. (1985). Child Care Arrangements and Fertility: An Analysis of Two-Earner Households. *Demography, 22*(4), 499-513.
Liefbroer, A. C., & Corijn, M. (1999). Who, What, Where and When? Specifying the Impact of Educational Attainment and Labour Force Participation on Family Formation. *European Journal of Population, 15*, 45-75.
Luci, A., & Thevenon, O. (2012). The impact of family policy packages on fertility trends in developed countries. *Ined: Documents de Travail*(174).
Mason, K. O., & Kuhlthau, K. (1992). The perceived impact of child care costs on women's labor supply and fertility. *Demography, 29*(4), 523-543.
Matysiak, A., & Vignoli, D. (2008). Fertility and Women's Employment: A Meta-analysis. *European Journal of Population, 24*, 363-384.
Matysiak, A., & Węziak-Białowolska, D. (2016). Country-Specific Conditions for Work and Family Reconciliation: An Attempt at Quantification. *European Journal of Population*, 1-36.
McDonald, P. (2000). Gender Equity in Theories of Fertility Transition. *Population and Development Review, 26*(3), 427-439.
Merla, L., & Deven, F. (2013). Belgium Country Note. In P. Moss (Ed.), *International Review of Leave Policies and Research 2013*.
Mills, M., Rindfuss, R. R., McDonald, P., & Te Velde, E. R. (2011). Why do people postpone parenthood? Reasons and social policy incentives. *Human Reproduction Update, 17*(6), 848-860.
Morel, N. (2007). From subsidiarity to "free choice": Child- and elderly-care policy reforms in France, Germany, Belgium and the Netherlands. *Social Policy & Administration, 41*(6), 618-647.
Neels, K. (2006). *Reproductive Strategies in Belgian Fertility, 1960-1990*.

Neyer, G. (2003). Family Policies and Low Fertility in Western Europe. *MPIDR Working Paper WP, WP 2003-021*.

Neyer, G., & Andersson, G. (2008). Consequences of Family Policies on Childbearing Behavior: Effects or Artifacts? *Population and Development Review, 34*(4), 699-724.

Ni Bhrolchain, M., & Beaujouan, E. (2012). Fertility Postponement is largely due to rising educational enrolment. *Population Studies, 66*(3), 311-327.

ONE. (2018). Office de la naissance et de l'enfance - Accueil petite enfance.

Plantenga, J., Scheele, A., Peeters, J., Rastrigina, O., Piscova, M., & Thévenon, O. (2013). Barcelona Targets Revisited. Brussels: European Parliament.

Population Council. (2006). Policies to Reconcile Labor Force Participation and Childbearing in the European Union. *Population and Development Review, 32*(2), 389-393.

Puur, A., Klesment, M., Rahnu, L., & Sakkeus, L. (2016). *Educational gradient in transition to second birth in Europe: differences related to societal context*. Paper presented at the European Population Conference (EAPS), Mainz, Germany.

Raz-Yurovich, L. (2014). A Transaction Cost Approach to Outsourcing by Households. *Population and Development Review, 40*(2), 293-309.

Rindfuss, R. R., & Brewster, K. L. (1996). Childrearing and Fertility. *Population and Development Review, 22*, 258-289.

Rindfuss, R. R., Guilkey, D. K., Morgan, S. P., & Kravdal, O. (2010). Child-Care Availability and Fertility in Norway. *Population and Development Review, 36*(4), 725-748.

Rindfuss, R. R., Guilkey, D. K., Morgan, S. P., Kravdal, O., & Guzzo, K. B. (2007). Child care availability and first-birth timing in Norway. *Demography, 44*, 345-372.

Ruokolainen, A., & Notkola, I.-L. (2002). Familial, situational and attitudinal determinants of third-birth intentions and their uncertainty. *Yearbook of Population Research in Finland, 38*, 179-206.

Shapiro, D., & Mott, F. L. (1994). Long-term employment and earnings of women in relation to employment behavior surrounding the first birth. *Journal of Human Resources, 29*, 248-276.

Sjöberg, O. (2004). The Role of Family Policy Institutions in Explaining Gender-Role Attitudes: A Comparative Multilevel Analysis of Thirteen Industrialized Countries. *Journal of European Social Policy*(14), 107-123.

Sobotka, T., Skirbekk, V., & Philipov, D. (2011). Economic Recession and Fertility in the Developed World. *Population and Development Review, 37*(2), 267-306.

Thevenon, O. (2008). Family policies in Europe: available databases and initial comparisons. *Vienna Yearbook of Population Research*, 165-177.

Thevenon, O. (2011). Does Fertility Respond to Work and Family-life Reconciliation policies in France? In N. Takayama & M. Werding (Eds.), *Fertility and Public Policy: How To Reverse the Trend of Declining Birth Rates*. Cambridge, London: MIT Press.

Thomese, F., & Liefbroer, A. (2013). Child Care and Child Births: the role of grandparents in the Netherlands. *Journal of Marriage and Family, 75*(2), 403-421.

Van Bavel, B., & Rozanska-Putek, J. (2010). Second birth rates across Europe: interactions between women's level of education and child care enrolment. *Vienna Yearbook of Population Research, 8*, 107-138.

Van de Kaa, D., & Lesthaeghe, R. (1986). *Bevolking, groei en krimp*: Van Loghum Slaterus.

Van Lancker, W., & Ghysels, J. (2012). Who benefits? The social distribution of subsidized childcare in Sweden and Flanders. *Acta Sociologica, 55*(2), 125-142.

Vande Gaer, E., Gijselinckx, C., & Hedebouw, G. (2013). *Het gebruik van opvang voor kinderen jonger dan 3 jaar in het Vlaamse Gewest.* Leuven.

Vandelannoote, D., Vanleenhove, P., Decoster, A., Ghysels, J., & Verbist, G. (2013). Maternal employment: the impact of triple rationing in childcare in Flanders. *KULeuven Center for Economic Studies: Discussion paper series DPS13.07*.

Vandenbroeck, M. (2006). Globalisation and privatisation: The impact on childcare policy and practice. *Working papers, 38, The Hague, The Netherlands: Bernard van Leer Foundation*.

Wood, J., & Neels, K. (2017). First a job, then a child? Subgroup variation in women's employment-fertility link. *Advances in Life Course Research, 33*, 38-52.

Wood, J., Neels, K., & Kil, T. (2014). The educational gradient of childlessness and cohort parity progression in 14 low fertility countries. *Demographic Research, 31*(46), 1365-1416.

Wood, J., Neels, K., Marynissen, L., & Kil, T. (2017). Differences de genre dans la participation au marche de l'emploi apres la constitution de famille. Impact des caracteristiques initiales du travail sur la sortie et la prise du conge parental dans les menages a deux revenus. *Revue Belge de Sécurité Sociale, 2*(5), 173-208.

Wood, J., Neels, K., & Vergauwen, J. (2016). Economic and Institutional Context and Second Births in Seven European Countries. *Population Research and Policy Review, 35*(3), 305-325. doi: 10.1007/s11113-016-9389-x

Wooldridge, J. M. (2002). *Econometric analysis of cross section and panel data*. Cambridge (MA), London: MIT Press.

**Figure 1**: Childcare coverage (number of places/population aged 0-3, %) in 588 Belgian municipalities in 2001-2004

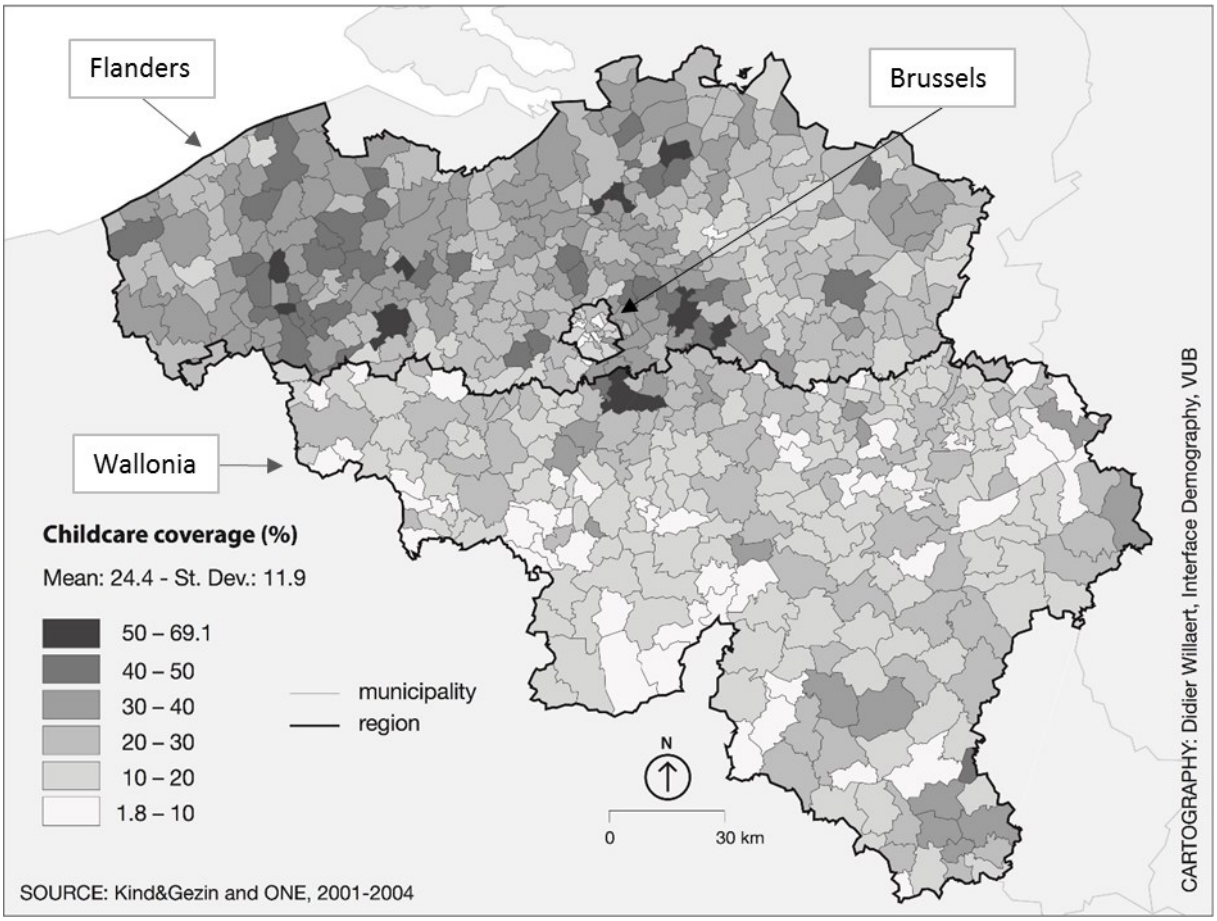
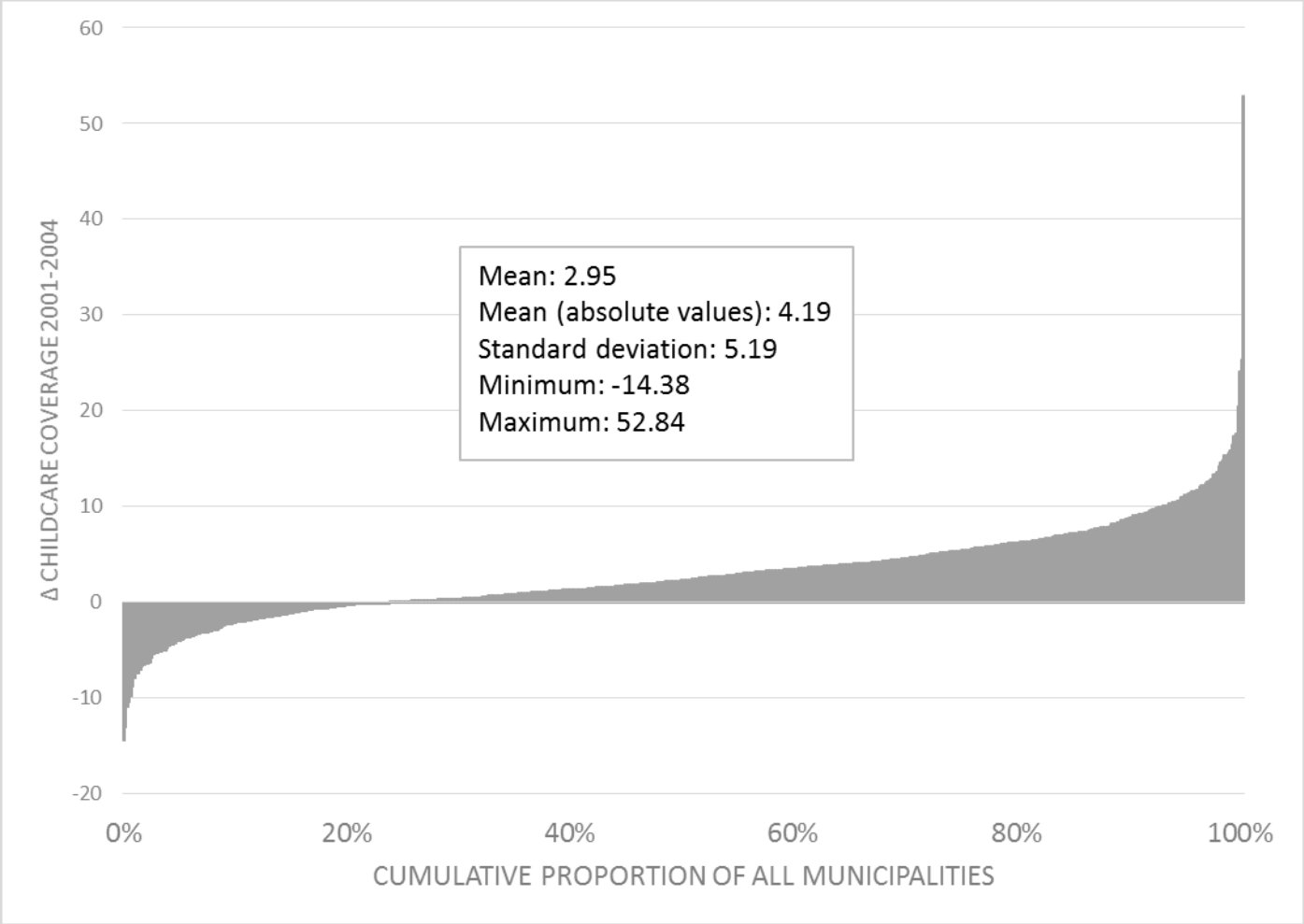
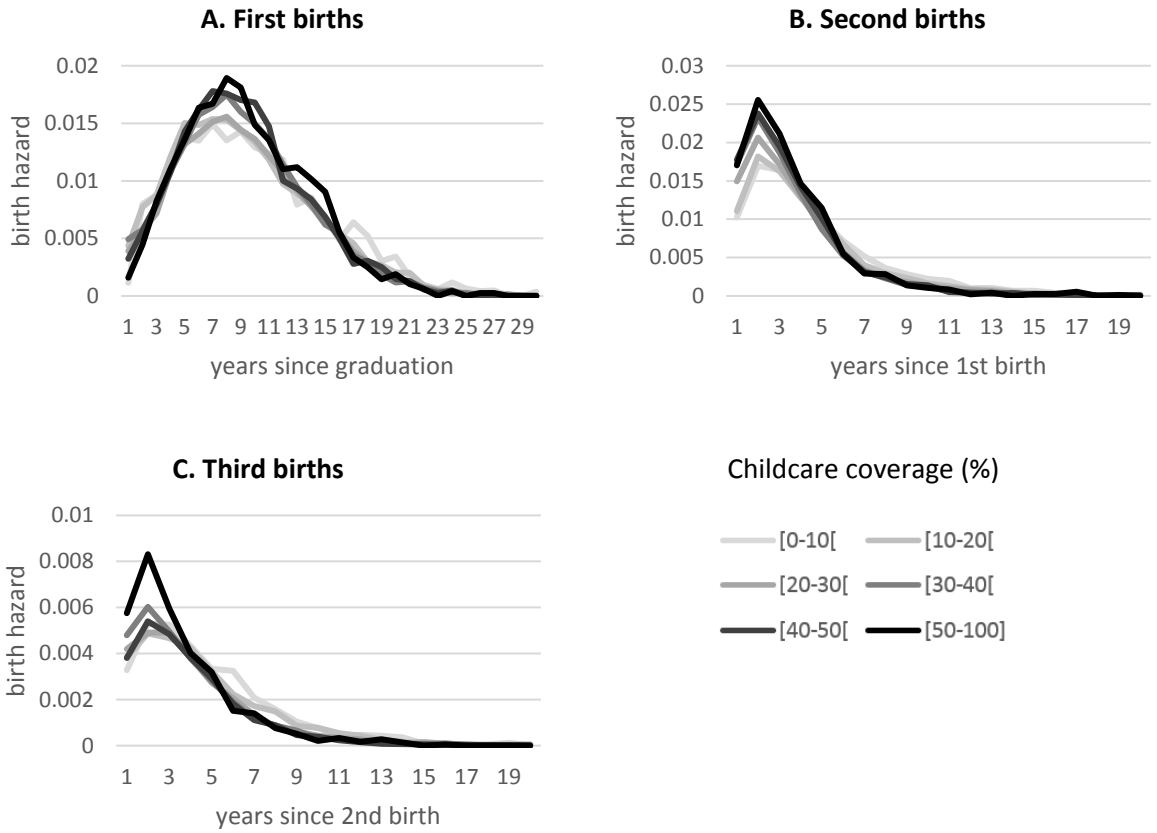


**Figure 2**: Evolution in childcare coverage within municipalities (number of places/population aged 0-3, %) in 588 Belgian municipalities in 2001-2004




*Source:K&G, ONE*

**Figure 3**: Observed first, second and third birth hazard functions by childcare coverage, Belgium 2001-2004.




*Source:2001 Belgian Census, Register, K&G, ONE*

**Table 1**: Exponentiated effects (odds-ratios) from Random-effects and Fixed-effects discrete-time event history models for first second and third births in Belgium, 2002-2005

| | **Model 1A Random-effects models** | | | | | | **Model 1B Fixed-effects models** | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 1st birth | | 2nd birth | | 3rd birth | | 1st birth | | 2nd birth | | 3rd birth | |
| | *e(b)* | *sig* | *e(b)* | *sig* | *e(b)* | *e(b)* | *sig* | *sig* | *e(b)* | *sig* | *e(b)* | *sig* |
| **Childcare Coverage** | | | | | | | | | | | | |
| *. Main effect + 10%* | 1.161 | *** | 1.051 | *** | 1.051 | *** | | | | | | |
| *. Main effect + 1%* | | | | | | | 1.108 | *** | 1.028 | *** | 1.021 | *** |
| | | | | | | *Model parameters* | | | | | | |
| Df. | 79 | | 67 | | 91 | | 660 | | 648 | | 672 | |
| -2LL | 478181.2 | | 376738.6 | | 191422.38 | | 523751 | | 375713.4 | | 190433.3 | |
| BIC | 479401.4 | | 377794.6 | | 192899.4 | | 485768 | | 385926.1 | | 201323.9 | |
| N Person-months | 5,096,609 | | 6,992,385 | | 11,191,065 | | 5,096,609 | | 6,992,385 | | 11,191,065 | |

*Significance levels: p < .050 (*), p < .010 (**), p < .001 (***) Source:2001 Belgian Census, Register, K&G, ONE*

**Table 2**: Aggregate-level fertility measures $SPPR_{1-3}$ and $SPPR_{1-3}$-based Total Fertility Rate, observed and under a 10% difference (random effects model) or 1% change (fixed effects model) in childcare coverage, Belgium 2002-2005

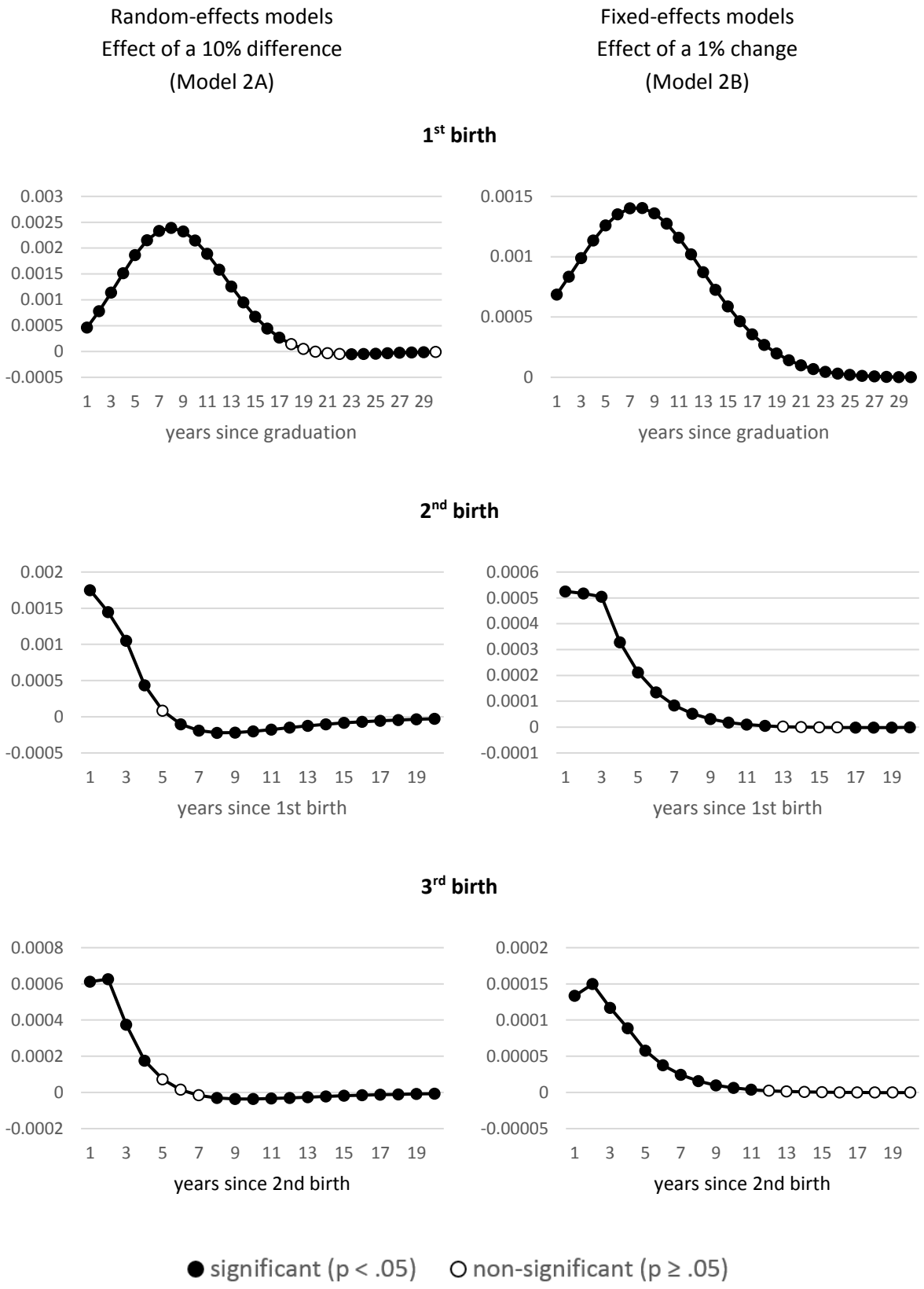
| | Observed Fertility[1] | + 10 % childcare coverage in random effects model | | | + 1 % childcare coverage in fixed effects model | | |
|---|---|---|---|---|---|---|---|
| | level | Level | Diff.[2] | Diff.[2] (%) | Level | Change[2] | Change[2] (%) |
| $SPPR_1$ | .895 | .921 | .026 | 2.91 | .915 | .02 | 2.23 |
| $SPPR_2$ | .680 | .692 | .012 | 1.76 | .689 | .009 | 1.32 |
| $SPPR_3$ | .288 | .301 | .013 | 4.51 | .293 | .005 | 1.74 |
| $SPPR_{1-3}$-based TFR | 1.679 | 1.750 | .071 | 4.23 | 1.730 | .051 | 3.04 |

*1: Observed fertility based on the average fertility schedule for our sample*
*2: Difference (random effects model) or change (fixed effects model) compared to observed fertility levels*

*Source:2001 Belgian Census, Register, K&G, ONE*

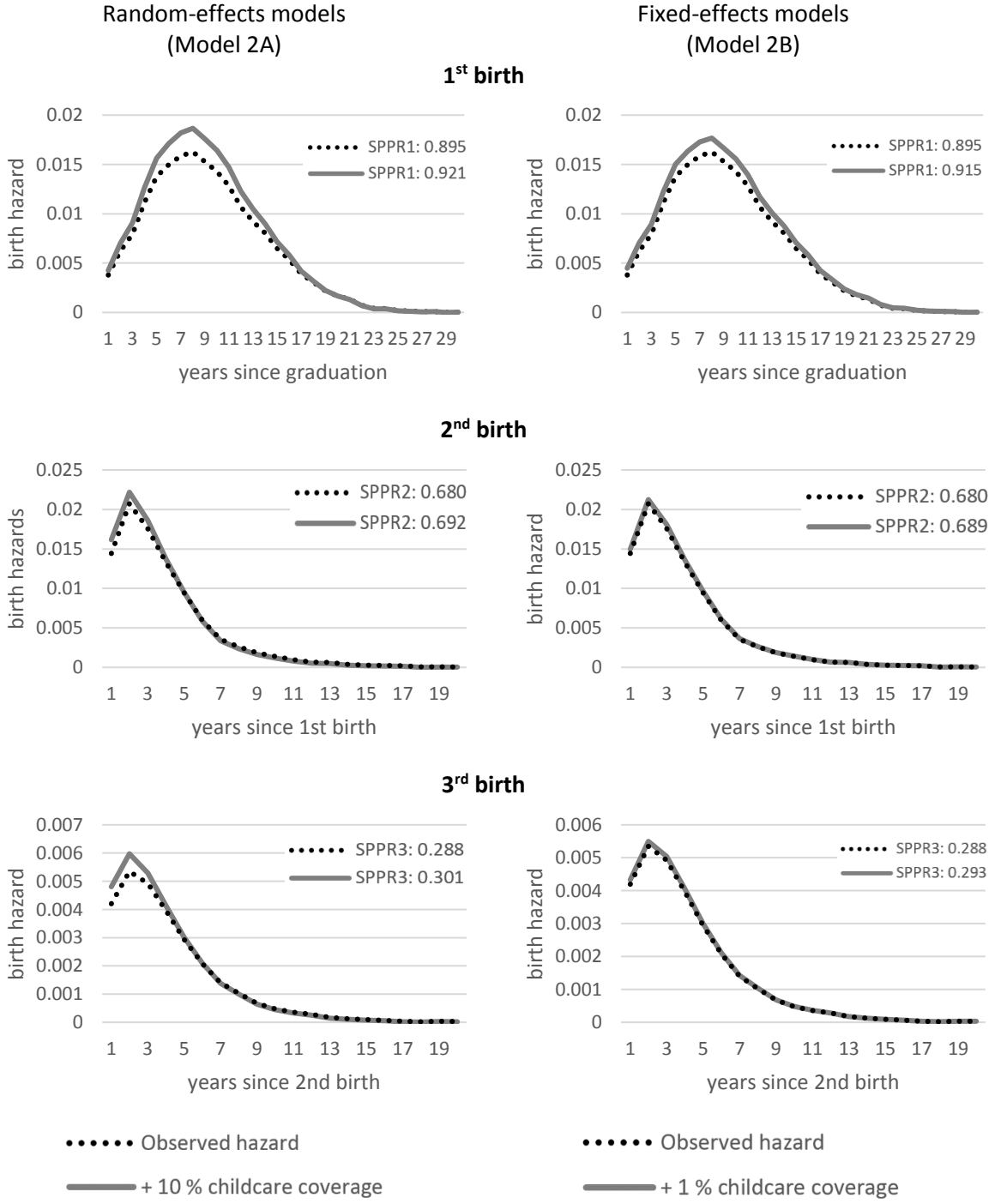
# APPENDIX

**Figure A1**: Marginal effects of childcare coverage on first, second and third birth hazards, Model 2, Belgium 2002-2005.




*Source:2001 Belgian Census, Register, K&G, ONE*

**Figure A2**: Observed birth schedules for Belgium versus the sum of observed birth schedules and marginal effects of childcare coverage, Model 2, Belgium 2002-2005.




*Source:2001 Belgian Census, Register, K&G, ONE*

**Table A1**: Exponentiated effects (odds-ratios) from Random-effects and Fixed-effects discrete-time event history models for first second and third births in Belgium, 2002-2005

| | Model 1A Random-effects models | | | | | | Model 1B Fixed-effects models | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 1st birth | | 2nd birth | | 3rd birth | | 1st birth | | 2nd birth | | 3rd birth | |
| | *e(b)* | *sig* | *e(b)* | *sig* | *e(b)* | *e(b)* | *sig* | *sig* | *e(b)* | *sig* | *e(b)* | *sig* |
| **Age male partner** | | | | | | | | | | | | |
| *. Under 25* | .763 | *** | .738 | *** | 1.164 | | .796 | *** | .742 | *** | 1.161 | |
| *. 25-30* | .964 | ** | .925 | *** | 1.007 | | .976 | | .926 | *** | 1.015 | |
| *. 30-34* | ref. | | ref. | | ref. | | ref. | | ref. | | ref. | |
| *. 35-39* | .829 | *** | .862 | *** | .875 | *** | .825 | *** | .860 | *** | .872 | *** |
| *. 40-45* | .625 | *** | .593 | *** | .637 | *** | .621 | *** | .590 | *** | .630 | *** |
| *. Over 45* | .393 | *** | .363 | *** | .470 | *** | .392 | *** | .360 | *** | .461 | *** |
| **Marital status** | | | | | | | | | | | | |
| *. Not married* | ref. | | ref. | | ref. | | ref. | | ref. | | ref. | |
| *. Married* | 3.178 | *** | 3.737 | *** | 1.881 | *** | 3.277 | *** | 3.780 | *** | 1.917 | *** |
| **Her origin group** | | | | | | | | | | | | |
| *. Belgium* | ref. | | ref. | | ref. | | ref. | | ref. | | ref. | |
| *. Neighbouring* | .996 | | 1.143 | *** | 1.072 | | .958 | | 1.113 | ** | 1.059 | |
| *. Southern Europe* | 1.031 | | 1.031 | | .886 | * | 1.026 | | 1.032 | | .887 | * |
| *. Other Europe* | .860 | ** | 1.061 | | 1.150 | | .858 | ** | 1.047 | | 1.123 | |
| *. Turkey, Morocco* | 1.278 | *** | 1.278 | *** | 1.447 | *** | 1.264 | *** | 1.269 | *** | 1.440 | *** |
| *. Non-Eur. highly developed* | .612 | ** | 1.220 | | 1.150 | | .613 | ** | 1.209 | | 1.115 | |
| *. other non-Eur.* | .982 | | 1.208 | *** | 1.427 | *** | .983 | | 1.196 | *** | 1.401 | *** |
| **His origin group** | | | | | | | | | | | | |
| *. Belgium* | ref. | | ref. | | ref. | | ref. | | ref. | | ref. | |
| *. Neighbouring* | 1.107 | *** | 1.080 | * | 1.227 | *** | 1.107 | *** | 1.055 | | 1.211 | *** |
| *. Southern Europe* | 1.153 | *** | 1.185 | *** | 1.077 | | 1.142 | *** | 1.190 | *** | 1.073 | |
| *. Other Europe* | 1.077 | | 1.204 | ** | 1.149 | | 1.059 | | 1.188 | * | 1.129 | |
| *. Turkey, Morocco* | 1.137 | ** | 1.234 | *** | 1.979 | *** | 1.118 | * | 1.218 | *** | 1.967 | *** |
| *. Non-Eur. highly developed* | .955 | | 1.312 | | 1.474 | | .947 | | 1.309 | | 1.438 | |
| *. other non-Eur.* | 1.050 | | 1.365 | *** | 1.790 | *** | 1.049 | | 1.352 | *** | 1.757 | *** |
| **Availability of non-employed adult in household** | | | | | | | | | | | | |
| *. no* | ref. | | ref. | | ref. | | ref. | | ref. | | ref. | |
| *. yes* | .988 | | .915 | | .913 | | .987 | | .918 | | .916 | |
| **Partners born in municipality of residence** | | | | | | | | | | | | |
| *. neither* | ref. | | ref. | | ref. | | ref. | | ref. | | ref. | |
| *. she* | .978 | | .941 | ** | .879 | *** | 1.002 | | .928 | ** | .856 | *** |
| *. he* | .961 | * | .934 | ** | .885 | *** | .986 | | .921 | *** | .860 | *** |
| *. both* | .979 | | .884 | *** | .758 | *** | 1.005 | | .869 | *** | .728 | *** |
| **Population density** (inhabitants per square kilometre) | | | | | | | | | | | | |
| *. < 250* | ref. | | ref. | | ref. | | . | | . | | . | |
| *. 250-499* | .852 | *** | .848 | *** | .867 | *** | . | | . | | . | |
| *. 500-749* | .800 | *** | .833 | *** | .860 | *** | . | | . | | . | |
| *. 750-1500* | .816 | *** | .838 | *** | .896 | ** | . | | . | | . | |
| *. > 1500* | .862 | *** | .914 | ** | 1.150 | *** | . | | . | | . | |
| **Female labour force participation** (childless women, age 18-49, 2001) | | | | | | | | | | | | |
| *. linear* | .985 | *** | .996 | | .985 | *** | . | | . | | . | |
| | | | | | | | *Model parameters* | | | | | |
| Df. | 79 | | 67 | | 91 | | 660 | | 648 | | 672 | |
| -2LL | 478181.2 | | 376738.6 | | 191422.38 | | 523751 | | 375713.4 | | 190433.3 | |
| BIC | 479401.4 | | 377794.6 | | 192899.4 | | 485768 | | 385926.1 | | 201323.9 | |
| N Person-months | 5,096,609 | | 6,992,385 | | 11,191,065 | | 5,096,609 | | 6,992,385 | | 11,191,065 | |

*Significance levels: p < .050 (*), p < .010 (**), p < .001 (***) Source:2001 Belgian Census, Register, K&G,ONE*
*Controlling for: age at entry in the risk set, exposure, education, education*exposure, work regime, work regime * exposure*