

Cash transfers, female labour force participation and fertility: from short to long run ^{*}

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Abstract

Developed countries face a demographic scenario of population decline driven by low fertility rates. As a response, governments are deploying family policies to increase the number of children. In this paper we propose a dynamic life-cycle model of fertility and female labour force participation, and use the short-term effects of a cash transfer policy from Spain, which the literature has cleanly identified, to calibrate its parameters. We retrieve the long-term effects via the model, and find that they are about half as large. Moreover, we find that an alternative policy consisting of childcare subsidisation yields only slightly smaller effects on fertility, but positive labour force participation effects. Finally, we explore how the coexistence of temporal and permanent contracts in Spain, which have different earnings profiles, affects fertility and interacts with cash transfers. The effects of this type of labour market are large, driven by increased returns to experience during crucial child-rearing years. However, the results on the effects of the cash transfers do not change in a scenario with a unique type of contract.

Keywords: cash transfers, fertility, female labour force participation, long-run, life-cycle

JEL Codes: J11, J13, J22

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1 Introduction

Historically, fertility rates have fallen with income and female labour force participation across countries and families. Economists initially stressed ideas like the quantity-quality trade-off of children (Becker, 1960; Becker and Lewis, 1973) and the role of the opportunity cost of women’s time (Butz and Ward, 1979) to explain these empirical regularities.¹ However, neither of the aforementioned two relationships hold anymore in developed countries: the first one has flattened and the second one is now slightly positive. The new economics of fertility stresses the importance of compatibility between career and family as a major driver of fertility, and sees it as crucial to consider fertility and labour supply decisions jointly (Doepke et al., 2022).

Moreover, in the developed world, the prevailing total fertility rates (TFR) have been well under the replacement level of 2.1 since the mid-1980s. This implies (in the absence of immigration) that the total population has already or will soon start to decline and that the share of old people will rise. The expected effects of these demographic changes include increasing burdens of social security systems (De Nardi et al., 1999), low interest rates (Krueger and Ludwig, 2007), reduced output growth and investment (Aksoy et al., 2019) and even the possibility of the economy entering a self-sustaining equilibrium path with declining population and living standards (Jones, 2020). Governments worldwide, worried about demographic pressures and their effects on the economy, are deploying policies to boost fertility rates and slow population decline and ageing.

However, the long-term effects of these policies on total fertility rates are not well understood. The empirical literature on family policies is most often able to cleanly identify short-term effects. These include both changes in timing (*tempo*) and total number of children (*quantum*). A policymaker that is worried about demographic trends is interested only in the latter. Still, there are good reasons to think that these policies may have an effect on long-run total fertility, chief among them the fact that there is a positive gap between desired and realised number of children in the developed world among recent cohorts of women. A crucial question is therefore how large the quantum effects of family policy are on fertility in the long run. To answer it, it is necessary to take into consideration how maternity and labour force participation interact.

This paper addresses these questions by proposing a dynamic life-cycle model featuring joint fertility and labour force participation decisions. To calibrate the model, we make use of a natural experiment involving a policy that gave cash transfers to women upon childbirth in Spain. The short-term fertility and labour force participation responses to this policy

¹For a survey on what Doepke et al. (2022) call first-generation models of fertility, see Hotz et al. (1993)

contain invaluable information regarding the sensitivity of such decisions to family income. Moreover, we match overall patterns in labour supply for groups of women with children of different ages, as well as moments reflecting the tempo and quantum of fertility decisions among Spanish women right before the implementation of the policy. We then use the model to quantify the long-term effect of cash transfers on the total number of children. Finally, we perform a series of policy experiments, implementing alternative policies to compare the fertility and labour force participation effects with those generated by cash transfers.

Our main result is that the long-term effects of cash transfers on fertility are only about half as big as the short-term effects, the main reason being that in the short term there are additional births among older cohorts of women that have only a few periods to adjust their decisions after the policy is announced. Moreover, we find that childcare subsidies for parents of children aged 0-3 that have the same present value as the cash transfers have an impact on fertility that is only slightly smaller than the latter, but increase labour force participation among mothers instead of reducing it (around 1 percentage point, while the unconditional transfers reduce it in a magnitude at least as large).

Furthermore, we explore the effects and interactions with the cash transfers arising from an important feature of Spanish labour markets: the coexistence of temporary and permanent contracts. Most workers start their careers in the former ones, which are akin to internships, and generate lower yearly earnings (partly due to intermittent unemployment). Therefore, the returns to experience are large in crucial child-rearing years, as women want to work, accumulate experience and increase the likelihood of transitioning to a permanent contract.

We find that that eliminating the dual nature of Spanish labour markets (i.e. implementing a unique contract with earnings between the temporary and the permanent ones) has a positive effect on fertility which is three times as large as that of the cash transfers. Finally, we explore the interaction between labour market duality and cash transfers. We find that the effects on fertility in the long run are very similar both in a scenario with dual labour markets and in one where there is a single contract (which is more representative of the situation in other developed countries like the United States). Taken together, our results highlight the importance that policy interventions that make labour market interruptions less costly (as in the elimination of duality) and less necessary (as in childcare subsidisation) have for fertility outcomes.

This paper contributes to two lines of research. First, it is related to the literature using dynamic life-cycle models of fertility and female labour force participation. While there is a large number of studies dealing with these issues in isolation, only a handful consider them as a joint decision². An even smaller number considers fertility decisions on a period-by-period

²Early papers studying fertility in a static context include [Becker \(1960\)](#), [Becker and Lewis \(1973\)](#) and

basis, as we do. [Moffitt \(1984\)](#), [Francesconi \(2002\)](#), [Da Rocha and Fuster \(2006\)](#), [Sheran \(2007\)](#), [Del Boca and Sauer \(2009\)](#), [Keane and Wolpin \(2010\)](#), [Adda et al. \(2017\)](#) and [Guner et al. \(2021\)](#) are some of them. One contribution to this literature is the way we deal with fertility preference heterogeneity. We provide evidence that the desired number of children, as retrieved from fertility surveys, changes very slowly in time, does not heavily depend on policy conditions, and provides a reasonable estimate of the lower bound for the potential fertility. Hence, we use desired fertility data to infer the distribution of preferences over the number of children. Moreover, we contrast the long-term effects of maternity on labour market outcomes in the model with those measured using the event-study methodology recently proposed by [Kleven et al. \(2019a\)](#).

This paper also contributes to the literature studying the impact of family policies on fertility and female labour force participation. There are broadly two branches in this body of research, one empirical and one structural (model-based). The empirical literature mainly uses difference-in-difference and regression discontinuity designs to exploit natural experiments created by policy implementation. In particular, [Milligan \(2005\)](#), [Lalive and Zweimüller \(2009\)](#), [Cohen et al. \(2013\)](#) and [González \(2013\)](#) study the impact of cash transfers³. The latter study provides the estimates for the short-term impact of this policy on fertility and mothers' labour supply that we use to calibrate our model.

On the other hand, the models used by the structural branch of the literature provide a laboratory to perform counterfactual experiments. This is important for three reasons. First, long-term impacts can be assessed this way. Second, the results from the empirical literature strongly suggest that the context in which family policies are implemented matters ([Cascio et al., 2015](#); [Olivetti and Petrongolo, 2017](#)). Thus, it is crucial to understand how they interact with one another and with labour market institutions. Policies can be implemented separately and jointly in the model, allowing researchers to better understand the interactions that arise. Third, it is possible to experiment with policies that have never been implemented, which means they cannot be tested empirically yet. Examples of this kind of work are [Erosa et al. \(2010\)](#), [Bick \(2016\)](#), [Adda et al. \(2017\)](#) and [Guner et al. \(2020\)](#). The main contribution of this paper is to connect the two branches. We believe that using the analytical tools from

[Willis \(1973\)](#). Dynamic models of fertility that kept labour force participation constant included [Ward and Butz \(1980\)](#), [Wolpin \(1984\)](#), [Rosenzweig and Schultz \(1985\)](#), [Cigno and Ermisch \(1989\)](#), [Heckman and Walker \(1990\)](#), [Blackburn et al. \(1993\)](#), [Arroyo and Zhang \(1997\)](#) and [Sommer \(2016\)](#). Conversely, [Heckman and Macurdy \(1980\)](#), [Blau and Robins \(1988\)](#), [Eckstein and Wolpin \(1989\)](#), [van der Klaauw \(1996\)](#), [Hyslop \(1999\)](#), [Attanasio et al. \(2008\)](#), [Keane and Sauer \(2009\)](#) and [Blundell et al. \(2016\)](#) study labour supply decisions in a dynamic framework, while taking fertility decisions as exogenous.

³Other family policies have been studied as well. [Bettendorf et al. \(2015\)](#), [Geyer et al. \(2015\)](#), [Givord and Marbot \(2015\)](#), [Haeck et al. \(2015\)](#), [Nollenberger and Rodríguez-Planas \(2015\)](#) and [Müller and Wrohlich \(2020\)](#) evaluate the effects of childcare subsidies, and [Asai \(2015\)](#) studies parental leaves.

the structural branch and calibrating the model’s parameters to carefully identified short-term effects is a powerful combination to understand better how family policies affect fertility and female labour supply.

The remainder of the paper is organised as follows. Section 2 discusses the setting and lays out some descriptive evidence. Section 3 describes the model. In section 4, we explain how we take the model to the data and show the calibration results. Section 5 presents the main quantitative experiments. Section 6 concludes.

2 Background and Descriptive Facts

In this section, we first briefly describe Spain’s demographics and compare them with those of similar countries. Then, we outline a natural experiment originated by a cash transfer policy that provides some key moments we use as part of our calibration targets, and describe the institutional context in which the policy was implemented. Moreover, we present some evidence on changes in labour market behaviour associated with motherhood. Finally, we present evidence on fertility preferences suggesting that there is a gap between the desired and realised number of children in Spain, and discuss the robustness of this finding. The facts presented in this section inform important modelling choices that are fully explained in section 3.

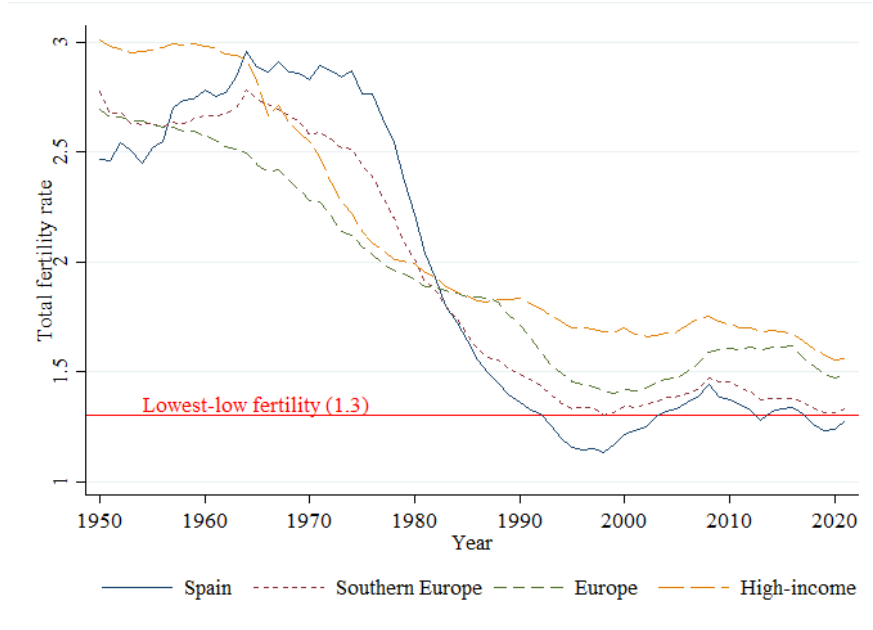
2.1 Spain’s demographics

In the early 1990s, Spain became one of the first countries in the world to attain what demographers call lowest-low fertility, i.e. a total fertility rate (TFR, from now onwards) below 1.3 (Kohler et al., 2002). A group of countries, mostly in Southern, Central and Eastern Europe, followed. Figure 1 shows how Spain experienced a relatively late baby boom in the 1960s and 1970s, during which its TFR increased above that of others in its geographic proximity and that of the rest of high-income countries⁴. However, it fell rapidly in the 1980s, dipping below the lowest-low fertility threshold for a good deal of the three decades between 1990 and 2022, and below that of its peers for the entirety of that period.

Moreover, since the early 1980s, most of the developed world has had a TFR below the replacement rate of 2.1. This means, in the absence of immigration, that population will eventually decline. Figure 2 shows the estimated and projected population in the countries for the same group of countries. In all of them, population peak is likely very close or has

⁴This Spanish baby-boom occurred later than that of other high-income countries, like the United States, where the baby-boom took place right after WWII and lasted until the mid-1960s.

Figure 1: Total fertility rate, selected countries (1950-2022)



Source: United Nations, Department of Economic and Social Affairs, Population Division (2022).

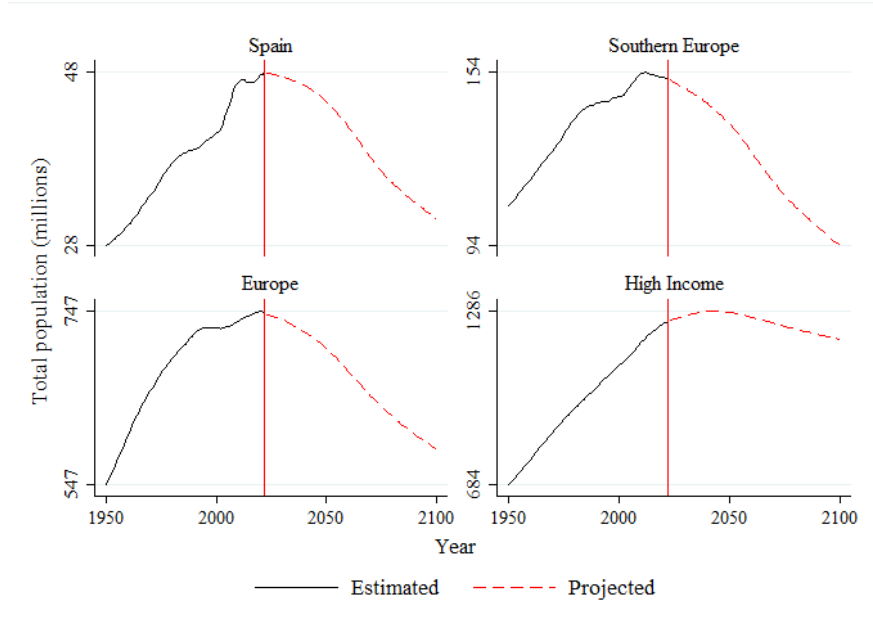
already passed. In the case of Spain, it is apparent that population would have already peaked if it were not for the large influx of immigrants (mainly from Latin America) it received in the 2000s. In fact, considering the very low fertility levels, in the absence of large-scale immigration, it is all but certain that most of the countries considered here will see population decline soon.

All said, Spain's demographic situation, while a bit more extreme than that of other similar countries, is still comparable in broad terms: TFR below replacement in all likelihood will lead to population ageing and decline.

2.2 The baby check

On July 3, 2007, the Spanish government unexpectedly announced the introduction of a one-time €2500 child benefit to be paid to mothers immediately after each birth. The policy came to be known in the Spanish media as “cheque bebé” (baby check, hereinafter). The projected ageing of the population is explicitly mentioned in the originating law as the motive behind the new benefit, with the implicit understanding that it would increase fertility and slow down that demographic trend. The mechanisms through which the government expected the baby check to stimulate fertility were also made explicit in the law: compensating the

Figure 2: Total population, selected countries (1950-2022)



Source: United Nations, Department of Economic and Social Affairs, Population Division (2022).

Note: Projections account for fertility, mortality and migration, and we use the medium scenario.

additional expenditures households incur upon the birth of a new child, and facilitating work-life balance. Lastly, the law mentions that the benefit was intended to maintain disposable income and welfare for low-income families that have a new child⁵.

The requirements to access the baby check were quite low, making it essentially universal. No income tests were stipulated, and the only real precondition was to have effectively resided in Spain for the two years before the birth⁶. This generated criticism to the benefit since its inception, as opposition parties argued that it was unfair that high-income families were receiving an identical benefit as low-income ones⁷. Nevertheless, the first baby checks were rolled out in November 2007, and over the next year, the government reported paying almost half a million of them, at a cost of approximately €1.2 billion, around 0.8% of the projected public expenditure for 2008⁸.

The sharp cut-off established for baby-check eligibility constitutes a clear natural exper-

⁵ LEY 35/2007, November 15, 2007.

⁶ Even this requirement was relaxed in 2009, making all women residing in Spain eligible for the benefit.

⁷ “El Congreso aprueba el cheque-bebé”. *El País*, October 18, 2007. https://elpais.com/elpais/2007/10/18/actualidad/1192695432_850215.html (in Spanish).

⁸ “Presupuestos Generales del Estado, 2008”. *Ministerio de Economía y Hacienda*, December 26, 2008.

iment, which was exploited by [González \(2013\)](#). Via a regression discontinuity analysis, she finds that mothers who received the benefit were 2-4 percentage points less likely to work twelve months after delivery. Moreover, she finds that the annual number of births increased by about 6 percent. These estimates provide invaluable information about the short-term sensitivity of households' fertility and female labour force participation decisions to unearned income.

The baby checks were phased out in 2010 as part of overall budget cuts implemented by the Spanish government after the 2008 financial crisis⁹. Nevertheless, the issue is not settled, neither in Spain nor elsewhere. OECD countries devote slightly more than 2% of GDP on average on family benefits public spending, with Spain actually devoting only around 1% of GDP ([OECD, 2022](#))¹⁰. Moreover, a new baby check was introduced by the regional government of Madrid in January 2022, this time geared towards young mothers (below 30 years old), means-tested (families with annual income below €30000) and much more generous (up to €14500, in €500 instalments)¹¹. The importance of assessing the effects of this kind of spending is therefore very much relevant.

2.3 Institutional background

The impact of family policies is mediated by the context in which they are implemented. Here, we discuss the institutional background in which the baby check was introduced in July 2007 in Spain.

Public childcare provision at this time in Spain was pretty typical for a high-income country. Despite mandatory schooling starting only at age 6, preschool between 9 am and 5 pm was offered to all parents of children aged 3 to 6 that requested it, and take-up was nearly universal. For children aged 0-3, however, the availability of places in public childcare centres was quite limited, and parents therefore mostly had to pay for it out-of-pocket. Using the *Encuesta de Presupuestos Familiares* (Family Budget Survey, EPF henceforth) from the Spanish Institute of Statistics, we estimate that families with a child aged between 0 and 3 years old using full-time childcare spent on average €2000 a year in 2007.

Taxation has been pointed at by the literature as one of the factors affecting female labour supply. In particular, joint taxation has been identified as a strong disincentive to labour force participation among married women ([LaLumia, 2008](#); [Bick and Fuchs-Schündeln, 2017](#)).

⁹Montero, Vicky, 2010. “José Luis Rodríguez Zapatero dice adiós al cheque-bebé, su medida estrella en política social”. *RTVE*, May 5. <https://www.rtve.es/noticias/20100512/331041.shtml> (in Spanish).

¹⁰Includes mainly child-related cash transfers, parental leaves and income support for sole parents, and services for families with children.

¹¹*Boletín Oficial de la Comunidad de Madrid* núm. 308, December 27, 2021.

Spain’s system, however, allows couples to minimize their tax liabilities by filing jointly or separately, so it did not constitute an additional barrier to female labour supply at this point.

One important aspect of the Spanish labour market that needs to be discussed here is duality, i.e. the coexistence of temporary and permanent contracts. Their main difference is firing costs, which are larger for permanent contracts. The result is that most young people start their careers as temporary workers. This causes job insecurity, loss of earnings caused by intermittent unemployment and increases the returns to labour market experience early on. In particular, temporary contracts increase the cost of career interruptions, as they make it less likely that the worker will be made permanent. There is plenty of evidence that temporary contracts cause women to postpone and reduce their fertility (see [Adserà \(2004\)](#), [de la Rica and Iza \(2005\)](#), [Auer and Danzer \(2016\)](#), [Lopes \(2020\)](#) and [Guner et al. \(2021\)](#)).

2.4 Desired and realised fertility

One of the reasons policies like the baby check may affect fertility is that in all high-income countries women on average would like to have more children than they do. In this section, we use data from the *Encuesta de Fecundidad* (Fertility Survey, EdF henceforth) of the Spanish Statistics Institute to discuss the robustness of this fact for Spain.

There are two waves of the EdF, from 1999 and 2018. Crucially, women were asked in both of them about their desired fertility. The structure of the question, which is identical across years, is as follows: first, women declare the number of children they have. If they are childless, they are asked whether they would like or would have liked children, and if so, how many. If they have children, they are asked whether the number they have coincides with the number they would like or would have liked to have. If it does not, they are asked how many children they would like or would have liked to have.

Table 1 shows the realised and desired fertility for women aged 40 to 44 in 2018, and the difference between them. We chose this age group because most of these women had already completed their fertility at this point, meaning that the difference with desired fertility is definitive. The main takeaway from this table is that there are many women whose completed fertility is lower than their ideal one. In particular, it shows that the fraction of childless women more than doubles the fraction of those who declare not wanting to have any children, while less than half the number of women that would have liked three or more children do. On average, the desired number of children for this cohort at this point in time was 2.04, while the actual number was 1.54, a gap of 0.5 children. This is very close to what is observed

on average across OECD countries¹².

We think that the existence of this gap is evidence that pronatalist policies have some margin to increase fertility. It can be reasonably argued that the answers given by people in surveys like the EdF do not represent policy-invariant preferences, and that people respond by taking into consideration restrictions and incentives given by their current context. However, a pronatalist policy would very likely not diminish desired fertility by itself. Therefore, the gap between it and realised fertility can be seen as a conservative estimate of the maximum possible effect of such a policy. Nevertheless, in what follows we present some evidence that desired fertility would probably not change significantly in response to policies that alter decisions in the margin.

Table 1: Fraction of women aged 40 to 44 by desired and realised number of children, Spain (2018)

| Number of children | Desired | Realised | Gap |
|--------------------|---------|----------|---------|
| 0 | 7.90% | 18.99% | -11.09% |
| 1 | 15.20% | 24.96% | -9.75% |
| 2 | 49.75% | 43.79% | 5.95% |
| 3 or more | 27.15% | 12.26% | 14.89% |

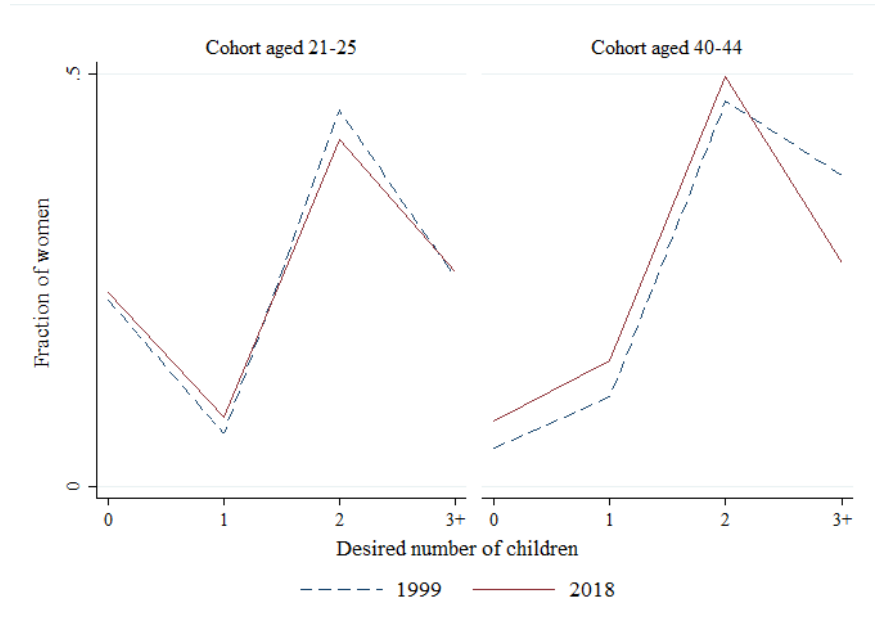
Source: *Encuesta de Fecundidad 2018, INE.*

An almost ideal way to test empirically whether the baby checks had any effect on desired fertility among the cohorts that were affected by it would be to use a RDD, similar to what [González \(2013\)](#) did for births and labour force participation. Unfortunately, we do not have data on desired fertility *right before* nor *right after* the announcement of the policy. We have two waves of the survey, which allows us to compare the same cohort’s responses at different ages and cohorts at the same age. Figure 3 shows cohorts aged 21 to 25 and 40 to 44 in 1999 and 2018. Consider first the cohort aged 21-25 in 1999 and 40-44 in 2018. These women were aged 29-33 in 2007, when the baby check was introduced, and therefore were affected by it. Their fertility preferences are represented by the dashed blue line in the left panel and by the solid red line in the right one. Notice that desired fertility experienced some changes between those years. In particular, the fraction of women that declared not wanting children in 1999 (when they were young, e.g. before most of them made any fertility decisions yet) is significantly smaller than the fraction in 2018 (after fertility was completed). Almost all of the difference is due to a larger proportion of women declaring a desired fertility of one child. On the other hand, the fractions wanting 2 and 3 or more children are similar across

¹²See OECD Family Database <https://www.oecd.org/els/family/database.htm>.

years. In other words, there is a discrepancy in the extensive margin, but not so much in the intensive margin. Two (non-mutually exclusive and non-exhaustive) hypotheses to explain this are that the baby checks changed the fertility preferences in such a particular way for this cohort, or that some women change their preferences as they get older. The left panel shows that the desired fertility is almost identical across cohorts when aged 21 to 25, and very similar across cohorts when aged 40 to 44 in 1999 and 2018. This is evidence in favour of the second hypothesis. In a nutshell, figure 3 points to fertility preferences changing somewhat along the life-cycle, but slowly across cohorts (which are exposed to different policies and economic conditions).

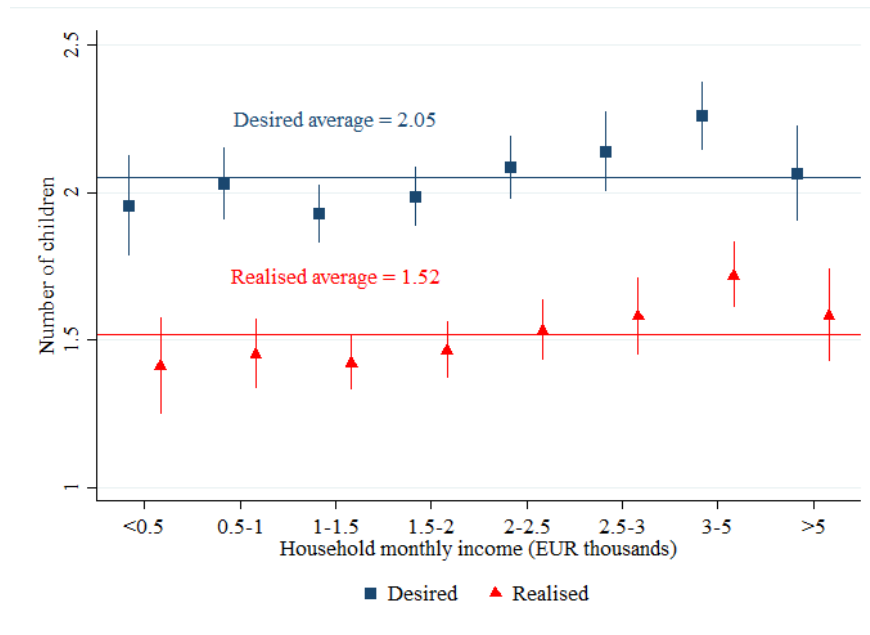
Figure 3: Fraction of women by desired number of children for selected cohorts, Spain (1999 and 2018)



Source: Authors' work with data from *Encuesta de Fecundidad 2009 and 2018*, INE.

Moreover, figure 4 shows the desired and realised fertility by household income for the same demographic as before. As discussed in the introduction, there's no clear relationship between fertility and income (if anything, there might be a small positive correlation, whereas in the past it was negative). Moreover, desired fertility also does not exhibit a clear pattern. Since it does not change a lot with differences in family income shown here, which go up to an order of 10, it seems reasonable to conclude that desired fertility barely reacts to a one-time cash transfer which is a fraction of that.

Figure 4: Desired and realised fertility by household monthly income among women aged 40-44, Spain (2018)



Source: Authors' work with data from *Encuesta de Fecundidad 2018*, INE.

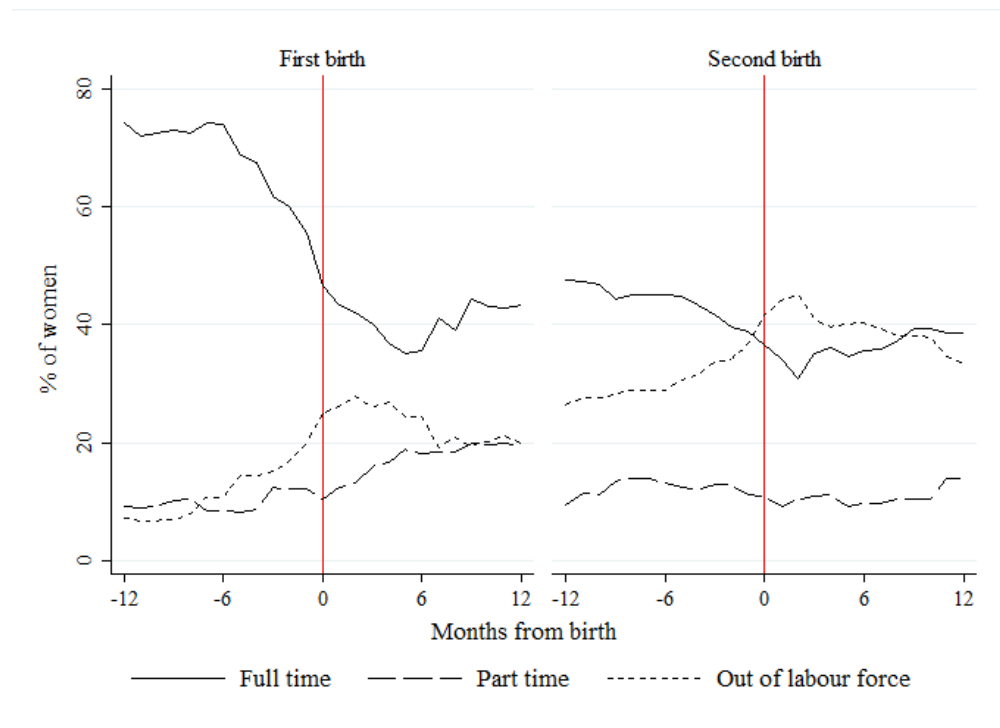
Note: Spikes represent 95% confidence intervals.

2.5 Labour force participation and motherhood

Motherhood is associated with important changes in women's labour market behaviour. In this subsection we use data from the *Encuesta de Condiciones de Vida* (Living Conditions Survey, ECV henceforth) of the Spanish Statistics Institute to provide some descriptive evidence for the magnitude of these changes among the cohorts of women affected by the baby check policy in the period immediately prior to its implementation. The ECV consists of a rotating panel, where households are interviewed for four consecutive years. There are 16000 households in the sample. Each year 4000 households leave, and 4000 new households replace them. We identified all births among women in the sample between 2004 (the first year of the sample) and July 2007 (that is, not eligible to receive the baby check). We kept all women for which there is labour force participation information for each of the 12 months preceding and succeeding the birth (24 months in total).

Figure 5 shows the fraction of women working full-time, part-time and not in the labour force in this period by parity (first and second births). There is a large drop in full-time participation that actually starts about six months before the birth of the first child. The difference in the rate twelve months before and after childbirth is more than 20 p.p. Con-

Figure 5: Female labour force participation around the first and second births in Spain, 2004-2007



Source: Author's work using *Encuesta de Condiciones de Vida 2005-2008*, INE.

versely, the part-time rate and the fraction of women that are out of the labour force increase by about 10 p.p. each¹³. By the time women give birth to their second child, full-time participation has barely changed compared to one year after their first one. However, the part-time rate has gone back approximately to its pre-children level, while the fraction out of the labour force has inched upwards by between 5 and 10 p.p. The second childbirth is associated with another drop in full-time participation rates, this time of slightly less than 10 p.p.. This is not compensated at all with part-time, and therefore almost fully translates in a commensurate increase in non-participation over the next 24 months.

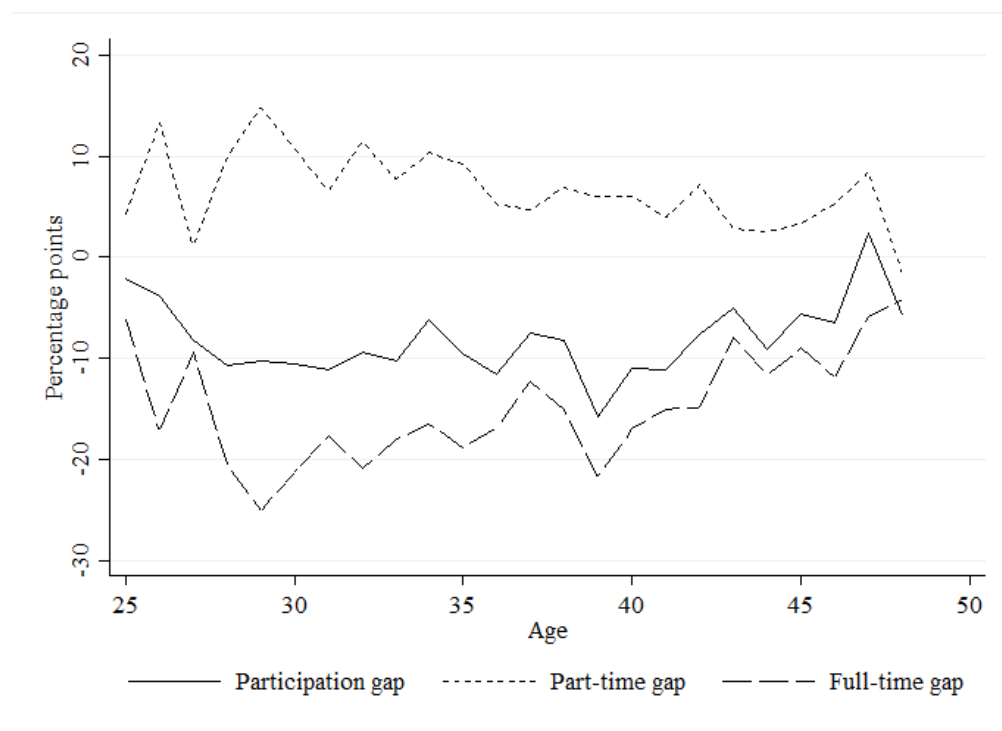
Taking stock, the first childbirth is associated with a large drop in full-time participation that is partially compensated with an increase in part-time, while the second birth is associated with a smaller drop that is, however, not compensated by an increase in part-time, and from a much lower baseline.

The previous analysis centres on the 24 months around childbirth and was performed on a balanced panel of women. To have an idea of the longer-term effects of maternity on labour market behaviour, we look at the cross-section of women present in the sample

¹³The fractions of women don't add up to 1 as some are also unemployed (not shown in the graph).

during the same period. Figure 6 plots the gap in participation, part-time and full-time rates between mothers and childless women by age. In general, mothers are much less likely to work full-time, somewhat more likely to work part-time and less likely to work overall. The participation gaps close as women age and presumably their children grow up and start requiring less time, thus allowing them to work more easily. However, they persist between ages 30 and 40. Therefore, maternity is associated not only with a sudden drop in participation the year around childbirth, but also with lower participation rates for several years afterwards.

Figure 6: Difference in labour force participation rates between mothers and childless women by age in Spain, 2004-2007



Source: Author's work using *Encuesta de Condiciones de Vida 2005-2008*, INE.

3 The model

We study fertility and labour force participation decisions over the life-cycle of married women. They derive utility from the number of children, consumption, and leisure. The model reflects the main trade-offs women face when making such decisions: each one has a target fertility level, but children negatively affect consumption and leisure time, so that desired fertility may not be achieved. The model captures dynamic career concerns as well.

There are various costs and benefits of becoming a mother at a younger and at an older age. Moreover, once a baby is born, women face a trade-off between being a working mother or taking a costly career break. Labour market participation interruptions are costly because experience is not accumulated. Moreover, most women start working life under a temporary contract. To increase the chances of converting it to a permanent one, they must work to accumulate experience. However, there are costs of being a working mother. Each woman decides how many children to have and when to have them, considering all these constraints. The details of the model are covered in the rest of this section.

3.1 Demographics

We model women's life choices between the ages of 25 and 52. A model period corresponds to one year. We denote model periods with j , hence age 25 corresponds to $j = 1$ and age 52 to $j = J = 28$. Prior to $j = 1$, each woman is exogenously matched with a spouse/partner that is the same age as her, and draws a desired number of children $N^* \in \{0, 1, 2, 3\}$. There are no marital transitions (no separations or divorces), and N^* remains constant for their whole life.

From ages 25 to 39, in each period women must decide whether to try to have an additional child or not. We denote this decision by $b \in \{0, 1\}$. If she decides to do so ($b = 1$), she gives birth to a baby next period with probability α_j , which decreases with age. Therefore, women can give birth between the ages of 26 and 40. Each woman can have a maximum of three children, and we denote the total number of children by $N \in \{0, 1, 2, 3\}$.

After being born, each child transitions stochastically between four stages of life, based on the different kinds of costs children impose on mothers as they age. The first two stages are newborn (below one year old, thus born in the current period) and baby (between ages 1 and 3). The only difference between them is that newborns may come with a baby check. Children in these two stages entail the same consumption, leisure and childcare costs. The third stage represents school-age (between ages 3 to 12). As children move to this stage, they continue to impose consumption and leisure costs, although quantitatively not identical to those imposed by newborns and babies. However, full-time childcare for school-age children is free for the parents (we assume universal public provision). Finally, those in the fourth stage are teenagers/young adults (ages 12 and up), which are not costly at all for mothers in terms of leisure, but represent a larger burden on consumption. The exact mathematical form all these costs take is discussed in the next subsection.

We denote by $\mathbf{n} = [n_0, n_1, n_2, n_3]$ the vector indicating the number of newborns, babies, school-age children and teenagers, respectively (hence $N = n_0 + n_1 + n_2 + n_3$). A newborn

becomes a baby with probability one after one period. A baby becomes a school-age child and a school-age child becomes a teen with probabilities $\lambda_1 = 1/2$ and $\lambda_2 = 1/11$, respectively. That is, in expectation, a child spends two years as a baby and eleven years in school-age. The teenage/young adult stage is absorbing: once a child reaches this stage, it remains there. We denote this structure by $\mathbf{n}' = \Lambda_j(\mathbf{n}, b)$ ¹⁴.

3.2 Preferences and constraints

Women are the sole decision-makers in the household. Alternatively, they hold all the bargaining power, and thus the household's preferences are perfectly aligned with hers. In each period, apart from their fertility decision, they must also decide their labour force participation $h \in \{0, 1/4, 1/2\}$, where $h = 0$ means no participation, and $h = 1/4$ and $h = 1/2$ stand for part-time and full-time work, respectively¹⁵.

In each period, utility is derived from consumption (c), leisure (l) and children (\mathbf{n}). The functional form for instantaneous utility is given by the sum of CRRA terms for the first two, and an additional term which is a function of the latter:

$$u(c, l, \mathbf{n}; N^*) = \frac{\left(\frac{c}{\psi(\mathbf{n})}\right)^{1-\gamma_c} - 1}{1 - \gamma_c} + \delta_l \frac{l^{1-\gamma_l} - 1}{1 - \gamma_l} + \Gamma[\mathbf{n}, h; N^*],$$

where:

$$\Gamma[\mathbf{n}, h; N^*] = -\delta_N(\mathbf{n}; N^*) \frac{\exp(j - \gamma_N)}{1 + \exp(j - \gamma_N)} |N(\mathbf{n}) - N^*| + \mathbb{1}_{\{N(\mathbf{n}) > 0\}} \zeta + \mathbb{1}_{\{h=1/2\}} \kappa(\mathbf{n}).$$

Children affect utility directly and indirectly. The first two terms of the above expression depend purely on the number of children, and are, therefore, the direct effect. First, women experience a utility penalty from the difference between current ($N(\mathbf{n})$) and desired or target number of children (N^*). This can be interpreted as craving for children among women that are still able to have them, and lifetime regret among those that cannot. Either way, it is the reason why women have children in the model.

The intensity of the penalty for not having the desired number of children depends on two terms. The first one, $\delta_N(\mathbf{n}; N^*)$ introduces a non-linearity:

¹⁴See appendix B for the exact functional form this structure takes.

¹⁵Out of the 24 hours a day has, we assume 8 are used for sleeping and personal care, leaving 16 hours to be split between work, childcare and leisure. A full-time job (8 hours per day) therefore represents $\frac{1}{2}$ of the time endowment, while a part-time one (4 hours per day) represents $\frac{1}{4}$ of it.

$$\delta_N(\mathbf{n}; N^*) = \delta_{N1} \left[1 - \delta_{N2} \mathbb{1}_{\{N(\mathbf{n})=2, N^*=3\}} \right],$$

i.e., we allow for the marginal disutility of not having a third desired child to be different from the marginal disutility of not having the first and the second.

The second one introduces age variation. In particular, it is a sigmoid function that increases with age and is asymptotic to 1. That is, younger women experience only a fraction of the utility penalty older women do, but this fraction increases with age. This can be interpreted in various ways. One is that it captures factors not included in the model that cause women to postpone fertility, such as housing and partner disposition. Another one is that it is a child-specific discount rate that decreases with age, as opportunities to have additional children diminish. Finally, there is a fixed (dis)utility of motherhood, ζ .

The functional forms described above allow the model to reflect aspects about the quantum and tempo of fertility. Anticipating the calibration, it should be apparent that ζ is crucial for the extensive margin of fertility (i.e. remaining childless or becoming a mother), δ_1 and δ_2 for the intensive margin (having 1, 2 or 3 children), and γ_N for the timing of births.

The third term in $\Gamma[\mathbf{n}, h; N^*]$ intends to capture the difficulties associated with working full-time while having kids. This includes schedule conflicts between full-time childcare and work schedules, and disutility from not spending time with children¹⁶. We allow this cost to vary on the age of the youngest child:

$$\kappa(\mathbf{n}) = \begin{cases} \kappa_1 & \text{if } n_0 + n_1 > 0 \\ \kappa_2 & \text{if } n_0 + n_1 = 0 \text{ and } n_2 > 0 \\ 0 & \text{if } n_0 + n_1 + n_2 = 0, \end{cases}$$

that is, if there is a newborn or baby present in the household, the utility cost of working full time is κ_1 , if there aren't any newborns or babies, but there are school-age children, it is κ_2 , and if there aren't any children other than teenagers, the cost is zero.

Children are costly in terms of consumption (non-childcare), and older children more so. To reflect this, we use the OECD equivalence scale to adjust per-capita consumption in the

¹⁶Childcare usually goes from 9 am to 5 pm, but a full-time job may not correspond to that schedule. The fact that parents with young children must drop off and pick up their children at determined hours themselves or have someone else do it introduces an additional cost to working full-time. Moreover, [Guner et al. \(2021\)](#) report on the unusual organization of the workday in Spain, with long lunch breaks that create split-shift schedules. For example, 50% of workers are still at work at 6 pm in Spain, compared to 20% in the UK.

household. It assigns a value of 1 to the household head, 0.7 to each adult and 0.5 to each child. Teenagers are counted as adults, hence:

$$\psi(\mathbf{n}) = 1.7 + 0.7n_3 + 0.5(n_0 + n_1 + n_2).$$

Leisure time is also negatively affected by children. We assume that children require a minimum amount of time from the mother that increases sub-linearly to reflect economies of scale in the childcare production function (e.g. it doesn't take double the amount of time to prepare food for two children than it takes to prepare food for one). Moreover, we allow the minimum time required by children to vary on the age of the youngest. In particular:

$$\xi(\mathbf{n}) = \begin{cases} \xi_1 \sqrt{n_0 + n_1 + n_2} & \text{if } n_0 + n_1 > 0 \\ \xi_2 \sqrt{n_2} & \text{if } n_0 + n_1 = 0 \text{ and } n_2 > 0 \\ 0 & \text{if } n_0 + n_1 + n_2 = 0, \end{cases}$$

e.g., if there are two children in the household, the cost can be different depending on the ages. If both are school-age, the leisure cost for the mother is $\xi_2 \sqrt{2}$, but if one is a baby, the cost is $\xi_1 \sqrt{2}$. This provides flexibility for the model to reflect the needs children of different ages may have. Notice once again that teenagers do not impose any leisure costs.

With this, we can define leisure time for the mother, which is the residual of a time endowment of one unit minus hours at work h minus time required by children:

$$l = 1 - h - \xi(\mathbf{n}).$$

Finally, we assume that women in the labour force with newborns and babies are required to purchase childcare time for the same amount of time she works. Resources available for consumption at the household level are therefore the household's net income I minus childcare costs:

$$c = I - \lambda 2h(n_1 + n_2),$$

where λ is the cost of full-time childcare (we assume part-time childcare costs half as much)¹⁷.

¹⁷It is possible that childcare costs do not increase linearly with the number of children (because of sibling

3.3 Income

We seek to model income in a way that captures income risk faced by households, accounts for labour market duality, is capable of reflecting the degree of (after-tax) income inequality (across households and genders, i.e. reflects the gender wage gap) and accounts for the returns of experience at different ages. To this end, we model spouses' (gross) income as depending on a couple of correlated, persistent stochastic shocks, the type of contract available (temporary or permanent), and accumulated experience.

The stochastic shocks for the woman and her partner are given by ϵ^f and ϵ^m , respectively. The household draws a couple of initial shocks $(\epsilon_1^f, \epsilon_1^m)$ from an exogenous joint distribution, that subsequently evolves over time following an AR(1) process:

$$\begin{aligned} \epsilon^{f'} &= \phi^f \epsilon^f + \nu^f \\ \epsilon^{m'} &= \phi^m \epsilon^m + \nu^m \end{aligned}, \quad \begin{bmatrix} \nu^f \\ \nu^m \end{bmatrix} \sim N \left(\begin{bmatrix} \mu^f = 0 \\ \mu^m = 0 \end{bmatrix}, \begin{bmatrix} \sigma_{\nu^f}^2 & \rho \\ \rho & \sigma_{\nu^m}^2 \end{bmatrix} \right).$$

Experience is accumulated over time by working. We assume the partners always work full time, and therefore their experience in period j is $j - 1$, $\forall j \in \{1, \dots, J\}$. For women, experience changes from one period to the next according to:

$$x' = \begin{cases} x & \text{if } h = 0 \\ x \text{ w.p. } 1 - \pi_z & \text{if } h = \frac{1}{4} \\ x + 1 \text{ w.p. } \pi_z & \text{if } h = \frac{1}{4} \\ x + 1 & \text{if } h = \frac{1}{2}, \end{cases}$$

that is, experience remains constant if she does not work, increases by one year if she works full time, and increases by one year with probability π_z if she works part-time. Notice that if $\pi_z = \frac{1}{2}$, half a year of experience is accumulated with one year of part-time work. However, we allow for $\pi_z \neq \frac{1}{2}$, possibly lower, to reflect a penalty of working part-time on experience accumulation. We denote this structure as $x' = \Pi_x(x, h)$.

We model labour market duality only for women. They draw an initial contract type $z_1 \in \{0, 1\}$, where 1 denotes permanent contracts and 0 temporary ones. In subsequent periods the probability of having an permanent contract depends on experience, the type of contract available in the previous period and whether she worked in the previous period:

discounts, etc.). However, we were not able to find good sources to determine the shape of the function.

$$z' = \begin{cases} 0 & \text{if } z = 0 \text{ and } h = 0 \\ 0 \text{ w.p. } 1 - \pi_z(x) & \text{if } z = 0 \text{ and } h > 0 \\ 1 \text{ w.p. } \pi_z(x) & \text{if } z = 0 \text{ and } h > 0 \\ 1 & \text{if } z = 1, \end{cases}$$

that is, if a woman has a temporary contract and works, she becomes permanent with probability $\pi_z(x)$ (which depends on her accumulated experience). If she has a temporary contract and does not work, she will still only have a temporary contract available for her next period. If she already has a permanent contract, she will also have one in the next period (it is an absorbing state). We denote this structure as $z' = \Pi_z(x, h, z)$.

Putting all the elements together, full-time (potential) log income for the woman is:

$$\ln(y^f) = \eta_0^f + \Delta\eta_0^f \mathbb{1}_{z=1} + \left(\eta_1^f + \Delta\eta_1^f \mathbb{1}_{\{z=1\}}\right) x + \left(\eta_2^f + \Delta\eta_2^f \mathbb{1}_{\{z=1\}}\right) x^2 + \epsilon^f,$$

that is, the type of contract changes the baseline income level and the returns to experience. In particular, if $\Delta\eta_0^f > 0$, $\Delta\eta_1^f > 0$, $\Delta\eta_2^f < 0$ and $\pi_z(x)$ is increasing in x , women have an additional reason to work as much as possible at the beginning of their career, i.e. to increase the likelihood of getting an open-ended contract, under which expected income is higher.

Log income for the husband is given by:

$$\ln(y^m) = \eta_0^m + \eta_1^m (j - 1) + \eta_2^m (j - 1)^2 + \epsilon^m.$$

The household's net income is the sum of the gross incomes of both partners minus tax liabilities:

$$I = y^m + 2hy^f \left(1 - \mathbb{1}_{\{h=\frac{1}{2}\}}\phi\right) - T(y^m, y^f, h),$$

where ϕ is an earnings penalty on part-time work, and $T(\cdot)$ is a tax liability function.

3.4 Timing, states, choice variables, and problem in recursive form

Upon entering the economy, women draw a fertility preference N^* , an initial contract type z_1 , and initial income shocks for them and their partners $(\epsilon_1^f, \epsilon_1^m)$. Then, they enter period 1 with no experience and no children, i.e. $x_1 = 0$ and $\mathbf{n}_1 = [0, 0, 0, 0]$.

From period 1 on, women observe their state vector $[\epsilon^f, \epsilon^m, z, x, \mathbf{n}]$, and choose their labour force participation h and whether or not to try to have an additional child next period $b \in \{0, 1\}$. This continues for every period, until they reach 3 children or age 39 ($j = 15$). After this happens, they cannot have any more children, and they choose only labour force participation in each period.

The dynamic problem women solve in period j is given by:

$$\begin{aligned}
V_j(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) = & \\
\max_{\substack{h \in \{0, \frac{1}{4}, \frac{1}{2}\} \\ b \in B_j(\mathbf{n})}} & u(c, l, \mathbf{n}; N^*) + \beta \mathbb{E} [V_{j+1}(\epsilon^{f'}, \epsilon^{m'}, z', x', \mathbf{n}'; N^*) \mid \epsilon^f, \epsilon^m, z, x, \mathbf{n}] \\
\text{s.t.} & \\
l = 1 - h - \xi(\mathbf{n}) & \\
c = I - \lambda 2h(n_1 + n_2) & \\
I = y^m + 2hy^f \left(1 - \mathbb{1}_{\{h=\frac{1}{2}\}}\phi\right) - T(y^m, y^f, h) & \\
\ln(y^f) = \eta_0^f + \Delta\eta_0^f \mathbb{1}_{z=1} + \left(\eta_1^f + \Delta\eta_1^f \mathbb{1}_{\{z=1\}}\right)x + \left(\eta_2^f + \Delta\eta_2^f \mathbb{1}_{\{z=1\}}\right)x^2 + \epsilon^f & \\
\ln(y^m) = \eta_0^m + \eta_1^m(j-1) + \eta_2^m(j-1)^2 + \epsilon^m & \\
\epsilon^{f'} = \phi^f \epsilon^f + \nu^f & \\
\epsilon^{m'} = \phi^m \epsilon^m + \nu^m & \\
\mathbf{n}' = \Lambda_j(\mathbf{n}, b) & \\
x' = \Pi_x(x, h) & \\
z' = \Pi_z(x, h, z), &
\end{aligned}$$

where the choice set for the birth decision is defined as:

$$B_j(\mathbf{n}) = \begin{cases} \{0, 1\} & \text{if } N(\mathbf{n}) < 3 \text{ and } j < 15 \\ \{0\} & \text{otherwise.} \end{cases}$$

4 Calibration

We calibrate the model in two steps. First, there is a set of parameters that we take from previous literature or estimate without solving yet the model. These include the parameters of the income process, the pregnancy success probabilities by age, the distribution of women over the number of desired children, and the cost of childcare. The second set of parameters is chosen jointly by targeting a set of data moments concerning average labour force participation rates for women without children and mothers with children of different ages, the tempo and quantum of fertility, and the short-term response to the sudden introduction of a cash transfer policy. For these, we rely on the Method of Simulated Moments.

4.1 Parameters chosen before solving the model

To estimate the parameters governing the household’s gross income, we turn to Spanish administrative data. In particular, we use the *Muestra Continua de Vidas Laborales* (MCVL, Continuous Working Life Sample), which is a 4% random sample of all affiliates to Social Security in Spain. The first step is to estimate the following Mincer regression separately for men, women with temporary contracts and women with permanent contracts:

$$\ln(y_{it}) = \beta_0^s + \beta_1^s x_{it} + \beta_2^s (x_{it})^2 + \Theta_{it} + \epsilon_{it},$$

where x_{it} is experience for individual i at time t , Θ_{it} is a vector of controls and $s = \{m, ft, fp\}$ stands for men, women under temporary and women under permanent contracts, respectively¹⁸. From the results of these regressions, we obtain the constant term and the returns to experience for the gross income equations for each of the three groups. We then take the residuals obtained from this estimation and regress them on their time lags at the individual level, to obtain the persistence parameters of the AR(1) process for the stochastic shocks, and the variance of the innovations. For the correlation coefficient between spousal shocks, we follow [Bick \(2016\)](#) and set a value of 0.25. The results of this procedure are shown in table 2.

For the numerical solution, we approximate the auto-regressive vector of stochastic shocks for the woman and her partner with a discrete-valued Markov chain using the method proposed by [Tauchen \(1986\)](#) and [Tauchen and Hussey \(1991\)](#). We use a 10 by 10 grid for the values of the shocks, where each point is calculated so that the income of the n-th point is the average income of the n-th decile for men and women at age 26. For the initial distribu-

¹⁸See the appendix for the exact specification.

tion of households over that grid, we identified all marriages without children and in which the woman was between 26 and 30 years old in the ECV between 2004 and 2007. Then, we computed hourly wages, and the deciles for each (separately for women and partners). Finally, we created a 10 by 10 matrix containing the fractions of couples by woman and partner decile¹⁹. This is the distribution from which the initial income shocks are drawn.

Table 2: Parametrisation of the income process

| | η_0^f | $\Delta\eta_0^f$ | η_1^f | $\Delta\eta_1^f$ | $\Delta\eta_2^f$ | η_2^f | ϕ^f | $\sigma_{\nu^f}^2$ |
|--------------|------------|------------------|------------|------------------|------------------|------------|----------|--------------------|
| Women | 6.348 | 0.806 | 0.048 | -0.007 | -0.003 | 0.001 | 0.906 | 0.184 |
| | η_0^m | | η_1^m | | η_2^m | | ϕ^m | $\sigma_{\nu^m}^2$ |
| Men | 7.093 | | 0.046 | | -0.001 | | 0.900 | 0.184 |

To estimate the probability of transitioning from a temporary to a permanent contract as a function of experience, we run a logit regression where the left-hand side variable is a dummy for being on a permanent contract, on experience and other controls. We then estimate the effect of one additional year of experience for the average individual, for each level of experience (from zero to 25 years). As a final step to obtain all the necessary parameters for household net income, we use the parametric effective tax functions estimated for Spain by [García-Miralles et al. \(2019\)](#) to account for tax liabilities and credits.

The probabilities of pregnancy success conditional on age α_j are estimated following [Sommer \(2016\)](#), who fits an exponential function to point estimates of infertility by age from the medical literature ([Trussell and Wilson, 1985](#)). They imply a success probability of 0.85 at age 25, which drops slowly to 0.77 at age 30, somewhat more rapidly to 0.65 at age 35 and then to 0.48 at age 39.²⁰

For the distribution from which women draw N^* , we use the fractions of women by desired number of children taken from the EdF reported in table 1. We are aware that these may be responsive to policy. However, in section 2, we discuss why it is unlikely that one like the baby check has a large effect on these answers, based on the small observed differences across average responses among people with different incomes in the cross-section, and the similarity of average responses across cohorts over time. In any case, we believe these answers capture relevant information about the heterogeneity in preferences among women in the population of interest and are useful in a calibration exercise.

¹⁹There is a small number of married, childless, non-working women in this age range, which we discard.

²⁰See appendix C for more details.

Using the Spanish Family Expenditures Survey, we estimated that a family with a child aged 0-3 using full-time childcare spent in 2007 on average around €2000. Therefore, we use this number as the cost per child per period for full-time, and half the amount as the cost for part-time childcare. Finally, for the subjective discount factor we use $\beta = 0.96$ (Kydland and Prescott, 1982).

4.2 Parameters chosen via the Method of Simulated Moments

The remaining 12 parameters are calibrated by matching 12 moments from the data. The targets can be divided in three groups. The first one comprises labour force participation rates for women with and without children. The second one encompasses average quantum and tempo of fertility. The third group involves the response to the cash transfers identified by González (2013).

We include part and full-time participation rates for three groups of women (6 targets): childless, mothers whose youngest child is 0-3 (newborn or baby) and mothers whose youngest child is 3-12 (school-age). The data moments are taken from the 2004-2007 ECV. For each woman, we have information on participation by month. We compute an average yearly participation rate, counting each month worked full-time as 1, each month worked part-time as $\frac{1}{2}$, and dividing by 12. Following Bick (2016), we create a yearly participation status variable that falls into one of our three categories: if the average participation rate was above 0.75, we count the woman as having worked full-time that year, if it is between 0.75 and 0.25 we count her as having worked part-time and if it falls below 0.25 we count them as being out of labour force. For childless women, we target the average part-time and full-time yearly rates between ages 25 and 51. For mothers, we target the average rates for all women with children of the respective age.

To compute the fertility targets, we use the EdF. In particular, we computed the fraction of women aged 40-44 in 2018 that had 0, 1, 2 and 3 children, and the average age at which this group had their first child. We use this sample because these women were in all likelihood very close to their completed fertility, which is the outcome of interest in the long run. These are 4 targets in total²¹.

Finally, González (2013) states in the conclusions of her paper that she finds that the number of births increased by 6 percent, and that mothers were 2-4 percentage points less likely to be working 12 months later as a result of the baby checks. We take these numbers as her preferred estimates and target a 6 percent increase in births and a 3 percentage points reduction in participation. These are the last two targets. To simulate the announcement of

²¹One of the fractions of women by the number of children is residual, and therefore there are really only three targets for quantum and one for the tempo of fertility.

the baby checks in the model, for each woman in each period we retrieve the policy function under the baseline scenario (no checks) and under an active policy scenario (baby checks announced). We store the decisions under both scenarios. We then calculate the average number of children born in a typical year without the policy, and the number of children that would be born with it, and compute the percent difference.

Although all parameters affect all model moments once we solve the model, some are more important than others for certain targets. The first two parameters, γ_c and γ_l govern how fast marginal utility from consumption and leisure falls, respectively. They are therefore important in determining how willing are women to substitute between consumption, leisure and children, and play an important role in the intensive margin of LFP decisions (part-time versus full-time) and for the fertility and LFP responses to the baby checks. The parameter on the age-varying weight on the fertility gap regulates how early the craving for children starts in the women’s life cycle, thus it is important for the average age at first birth. The weight on leisure δ_l plays an important role for extensive margin LFP decisions. For the distribution of women by number of children, δ_{N1} , δ_{N2} and ζ are crucial. The first one increases the penalty for not achieving the desired number of children, and therefore when it is higher every woman is more likely to be closer to her desired number of children. The second one diminishes the marginal penalty of not having a third child, and is thus important for the fraction of women that end up having 3 children. The last one is an additional utility from just being a mom, and plays an important role in the extensive margin of fertility, that is, the decision between remaining childless or having children. The next four parameters, ξ_1 , ξ_2 , κ_1 and κ_2 are relevant for the LFP decisions of mothers with children at different ages. The first two affect the extensive margin, while the last two the intensive one (by making full-time work costly). Finally, the part-time earnings penalty ϕ evidently has a direct impact on the likelihood of part-time work.

Table 3 shows calibrated parameters, with a description of their role in the utility function. Consumption utility is almost logarithmic, with γ_c very close to 1. The penalty for not having a third child is 27% lower than the penalty for not having the first two, while there is additional utility from the first child as given by the fixed utility from motherhood ζ being positive. This can be interpreted as the utility cost of remaining childless being larger than the cost of having children but fewer than desired, which seems reasonable. The time cost with younger children is larger, while the extra cost of full-time work is very similar independently of the age of the children. Finally, the model implies that working part-time entails a 20% penalty on hourly earnings compared to working full-time.

Table 4 shows the model’s outcomes versus the data targets. In general, we achieve our objective of reproducing closely the data targets with the calibrated model. In particular, it

Table 3: Parameters calibrated with the method of simulated moments

| Parameter | Description | Value |
|---------------|---|--------|
| γ_c | Curvature of consumption | 0.985 |
| γ_l | Curvature of leisure | 0.151 |
| γ_N | Age-varying weight on fertility gap | 26.500 |
| δ_l | Weight on leisure | 0.832 |
| δ_{N1} | Base weight on fertility gap | 0.364 |
| δ_{N2} | Weight on fertility gap, 3rd child | 0.270 |
| ζ | Fixed utility of motherhood | 0.115 |
| ξ_1 | Time cost, youngest child 0-3 | 0.349 |
| ξ_2 | Time cost, youngest child 3-12 | 0.211 |
| κ_1 | Cost of full-time work, youngest child 0-3 | -0.030 |
| κ_2 | Cost of full-time work, youngest child 3-12 | -0.040 |
| ϕ | Part-time earnings penalty | 0.200 |

is important that we were able to get the correct magnitude of the effect of the baby checks on fertility. We fall slightly short on the magnitude of the effect on LFP, but we are not too far from the lower range of 2 percentage points drop in [González \(2013\)](#). The effect is nevertheless small, and the model is qualitatively close.

4.3 Model evaluation: non-targeted moments

Here we discuss the model’s results for a set of non-targeted moments. First, we would like to know how well the model replicates the long-term costs that motherhood imposes on women. [Kleven et al. \(2019a\)](#) propose an event-study specification around the birth of the first child to measure the effect of children on the labour market outcomes of the parents. This specification, originally applied on Danish data, has since been used by [Kleven et al. \(2019b\)](#) for Sweden, Germany, Austria, the UK and the US. More importantly for us, [de Quinto et al. \(2021\)](#) did it for Spain, using the MCVL (which is the same data that we use to estimate the parameters of our income process). They assess the impact of the first child on gross earnings, days of work, probability of part-time employment and probability of being on a temporary contract. The first and the last two of these outcomes have direct counterparts in our model. We implement the event-study specification on our simulated data. The results, together with the original estimates by [de Quinto et al. \(2021\)](#), are shown in figure 7.

Qualitatively, the model displays adequate behaviour in all three outcomes: the earnings and part-time penalties increase rapidly the first two years after giving birth to the first child, while the temporary penalty is close to zero at the beginning and then increases slowly with

Table 4: The model vs. the data, baseline calibration targets

| Moment | Model | Data | Difference |
|---|--------|--------|------------|
| <i>Labour force participation:</i> | | | |
| <i>Childless women:</i> | | | |
| Part-time rate | 0.178 | 0.194 | -0.016 |
| Full-time rate | 0.690 | 0.717 | -0.027 |
| <i>Mothers, youngest child 0-3:</i> | | | |
| Part-time rate | 0.250 | 0.276 | -0.026 |
| Full-time rate | 0.486 | 0.537 | -0.051 |
| <i>Mothers, youngest child 3-12:</i> | | | |
| Part-time rate | 0.252 | 0.276 | -0.025 |
| Full-time rate | 0.536 | 0.583 | -0.046 |
| <i>Fertility:</i> | | | |
| <i>Share of women with:</i> | | | |
| 0 children | 0.171 | 0.190 | -0.019 |
| 1 child | 0.238 | 0.250 | -0.012 |
| 2 children | 0.460 | 0.438 | 0.022 |
| 3 children | 0.131 | 0.123 | 0.008 |
| Average age at first birth | 29.246 | 29.300 | -0.054 |
| <i>Effects of cash transfers on:</i> | | | |
| Annual number of births | 0.061 | 0.060 | 0.001 |
| Mother's LFP over the first year | -0.012 | -0.030 | 0.018 |

Note: The effect of the cash transfers is the number of births over the next year after the introduction of the policy. The effect on mothers' LFP over the first year is the difference between participation rates. Data points are the ones reported by [González \(2013\)](#) in the conclusions of her paper, which we take to be her preferred ones: 6 percent increase in the annual number of births, mothers 2-4 percentage points less likely to be working 12 months later.

age. Moreover, the three outcomes show a long-term penalty, i.e. even 10 years after giving birth the gap is still present.

The reason why earnings and part-time probability increase immediately after the first birth is straightforward: after having their first child, many women switch to part-time employment or drop out of the labour force, which means they earn less or nothing at all (the earnings child penalty is estimated unconditional on employment status). The reason behind the shape of the penalty on the probability of being on a temporary contract is more subtle: women on permanent contracts are more likely to have children (in the data you can see it as a negative penalty for the first two years, in the model, there is only a very small

effect the period the child is born). However, over time those that had children but were on temporary contracts convert them at a much lower rate than those who didn't have children, because many of the former take career breaks or switch to part-time.

Quantitatively, the model's long-term part-time and temporary contract probability penalties are quite close to the data counterparts. The main discrepancies occur in the first 5 years of the part-time and the last 6 of the earnings penalties. We think this is due to the fact that there are a number of factors that the model does not feature that affect earnings among mothers, including occupational choice and loss of skills (Adda et al., 2017). During the first years after birth, the model gets the right earnings by overestimating the part-time penalty. After the fifth year, the part-time penalty falls very close to the one from the data. The earnings penalty reflects it by diverging from the one from the data.

While we target average labour force participation by childless women and women with children of different ages, we would also like to have an idea of how well the model reproduces the participation gaps that exist between childless women and mothers at different ages. Figure 8 shows this, along with the data counterparts (which are the same series we showed in figure 6). The model reflects very well the general patterns: negative overall participation and full-time gaps and positive part-time gaps that are the largest around age 30 and close gradually as women age.

Finally, although we target the average age at first birth, it is interesting to compare how its distribution looks with respect to the data. This is shown in figure 9. Again, qualitatively the model reproduces the main features: the fraction of women having their first child is positive for every age between 26 and 40, with the bulk of women having it in their late 20s and early 30s, and the fraction of first-time mothers decreasing slowly after that. The main discrepancies occur with women aged 26 and with women in their mid-thirties. However, the model displays a good amount of time variability in this dimension.

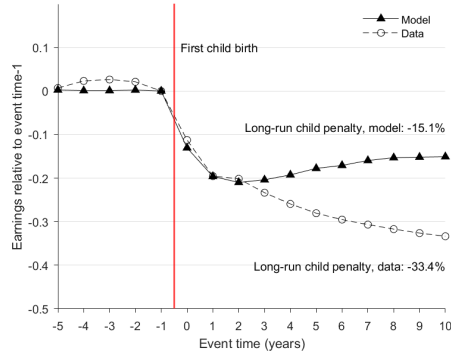
Overall, the fit of the model along non-targeted moments is satisfactory, and it seems to be an adequate setting to perform the quantitative experiments necessary to answer the main research questions considered in this paper.

5 Quantitative experiments

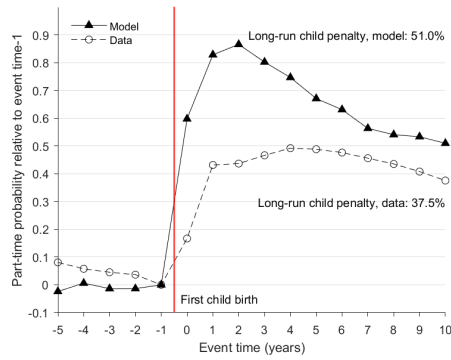
One of the main advantages of having a structural model like the one we propose in this paper is that it allows us to perform quantitative experiments consisting of counterfactual simulations. In this section, we present the results of three such experiments. For the first one, we assess the long-term effects of cash transfers on fertility, by simulating the life-cycle of women that are eligible for the policy for the entire duration of their lives. Then, we analyse

Figure 7: Child penalties in the model and the data

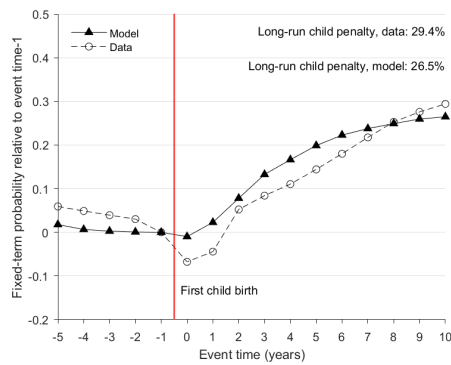
(a) Earnings



(b) Part-time work



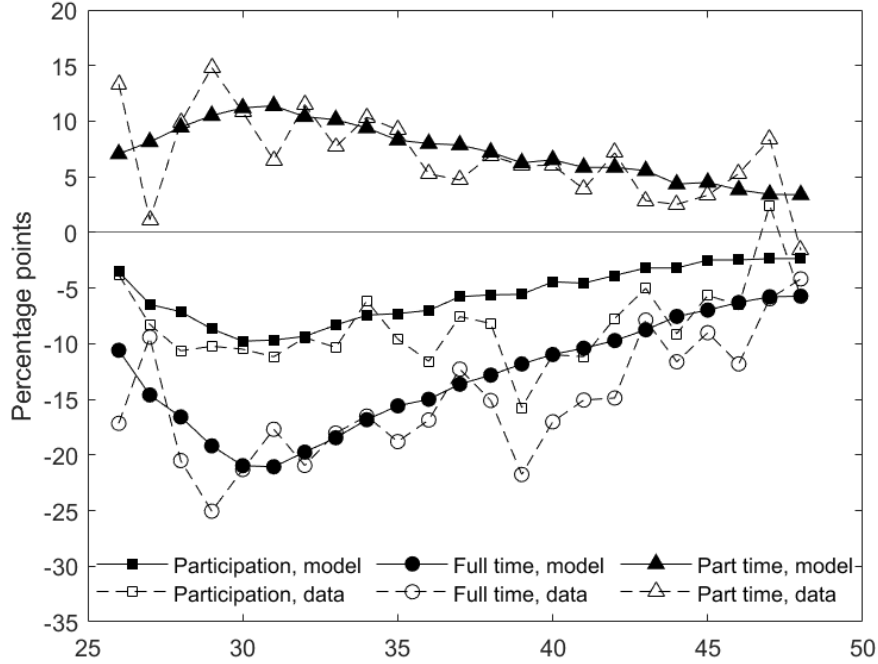
(c) Fixed-term contract



Source: Source: author's work with and data from [de Quinto et al. \(2021\)](#).

Note: The effects on (gross) earnings are estimated unconditional on employment status. The effects on part-time and fixed-term contracts are estimated conditional on working.

Figure 8: Labour force participation gap between mothers and childless women by age in the model



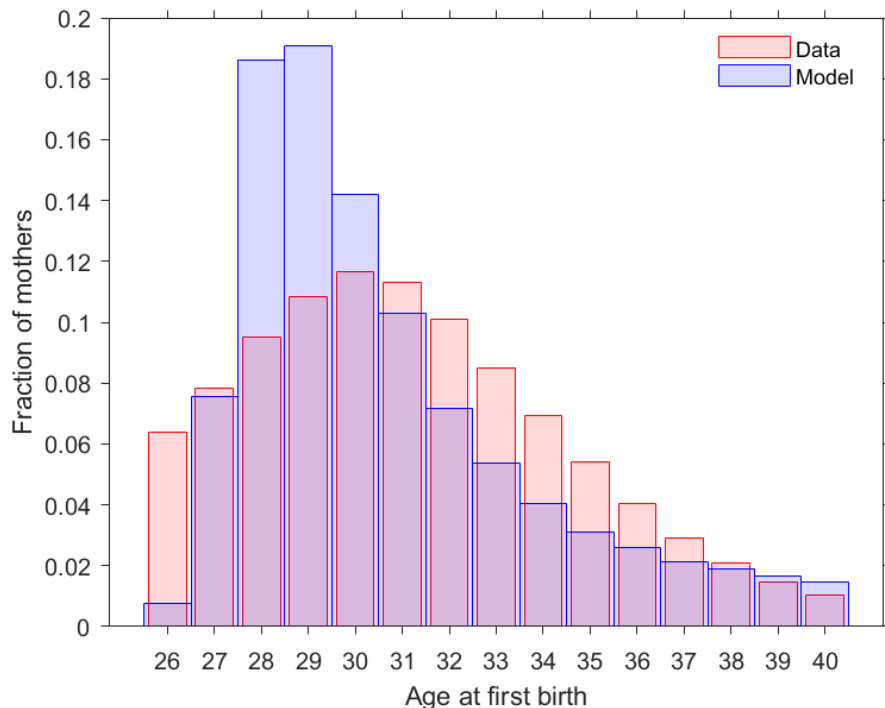
Source: Author's work and data from *Encuesta de Condiciones de Vida, 2004-2007*, INE.

the effects of an alternative policy, consisting of the subsidisation of childcare for mothers of children aged 0-3, which is the one group for which there was no public universal coverage at the time the baby check was introduced. Finally, we explore the role that the duality of the Spanish labour markets play on fertility, its interplay with labour force participation, and the effects of the cash transfers themselves.

5.1 Short and long-term effects of the cash transfer

The main objective of this paper is to provide an estimate of the effects of a cash transfer policy on fertility in the long run. We are therefore interested on the *completed fertility rate* (CFR, the average total number of children women have) in the long-run. To obtain this number, we simulate the life-cycle fertility and labour force participation decisions of women using our model, with every one of them being eligible to receive the baby check in every period of their life, and having full knowledge of the fact. Then, we can compare the CFR of a cohort of women in such a situation with the CFR in the baseline scenario, with no cash transfers.

Figure 9: Distribution of mothers by age of first birth, model and data



Source: author's work and data from *Encuesta de Fecundidad 2018*, INE.

In our model baseline, the CFR was 1.553 children per woman. In our counterfactual exercise in which each woman is eligible to receive the baby check in every period, the CFR goes up to 1.599. This is a 2.95% difference, just short of half of the 6% effect in the short run.

To understand why the long-term effects are different from those in the short term, it is useful to think about the effect of the policy on women of different ages, and to think about children as durable goods. After a certain age (around 30), most women want to have them as soon as possible (because they provide a utility flow). The exact timing depends on individual conditions. However, if a new policy that makes it more favourable to have children (at any age) is introduced, the likelihood that a woman has one in the following period goes up for all of them. However, for younger women, the likelihood that they have them later may decrease. The profile of the crude birth rate (CBR, the fraction of women that have a baby in a certain period of time) by age therefore may look very different for a cross-section of women at the time the policy is introduced compared to a cross-section of women a few years later. The effect of the baby check on fertility found by [González \(2013\)](#) is on the total number of births in a short window of time after the policy was implemented.

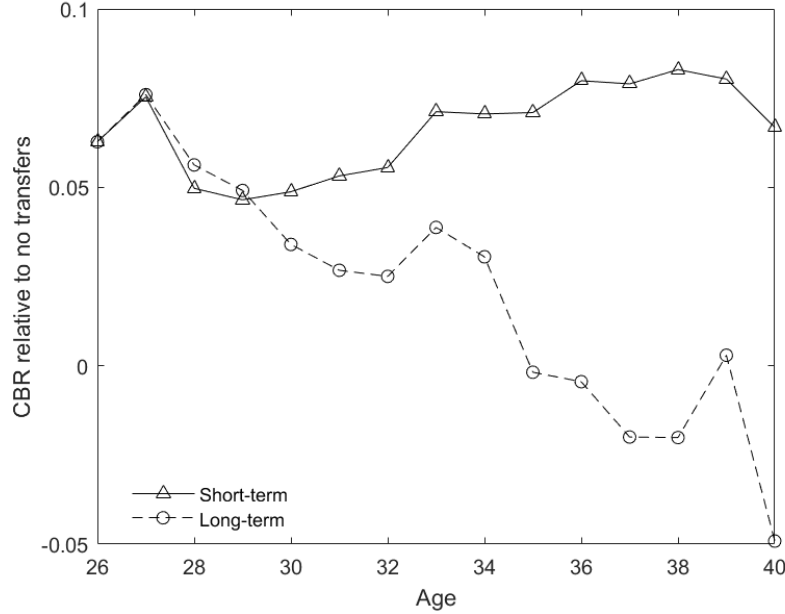
This effect depends directly on the former CBR profile. However, in the long-run, the CFR depends on the latter.

Figure 10 shows the effect of the cash transfer on the CBR by age in the short and the long-run. The short-term effect is the CBR by age of the cross-section of women living at the time of the policy relative to the baseline CBR by age. That is, it shows the average fertility response for women of every age between 26 and 40 the period following the policy announcement. The long-run effect is the CBR by age for women that were eligible for the baby check for all of their reproductive lives, i.e. for ages 25 to 39. Notice that the two lines are identical for young women. This is because these women face identical horizons. However, the two lines start to diverge in the 30s. The short-term effect remains positive, and has a soft U-shape. The reason is that cash transfers have a larger effect on young women, who are poorer, and on older women, for whom the transfer represents an incentive to take one of the few last opportunities to close the gap between their realised and desired number of children. Women in their late 20s and early 30s were having children anyways, so the effect on them is smaller. The long-term effect starts to fall, and even becomes negative in the late 30s. That is because young women had time to plan the births of the children they wanted before, and there are no older women who are caught by surprise by the policy anymore.

To further understand how the effect of cash transfers on births changes over time, we simulate fifteen cohorts of women of different ages when the baby checks are announced. In year -1, the policy is announced. Then, in year 0, the overall crude birth rate depends on the responses of women aged 25-39 that were caught by surprise by the policy in the previous period and changed their decisions accordingly. In year 0, women aged 26-39 make fertility decisions conditional on their previous period choices, which affect the next period. The average decisions of each cohort in period 0 may be different from the average decisions of the previous cohort in period -1 (when they were the same age). This is because some women in the newer cohort had an additional child, and that lowers the likelihood that they will have one during this period. Figure 11 shows the crude birth rate by year after the implementation of the policy. The CBR in the period right after the implementation goes up by around 6%, but then it gradually falls to around 3%, as the CBR profile by age converges to the long-term one shown in figure 10.

By observing the shape of the long-term effects of cash transfers by age one concludes that timing must also change. In figure 12 we have plotted the fraction of women by age at first birth in the baseline and in our long-term counterfactual exercise. As expected, the distribution under the policy has more mass at younger and lower mass at older ages. In particular, there are more first-time mothers in their late 20s and fewer in their late 30s when

Figure 10: Effect of the cash transfers on the crude birth rate (CBR) by age



Note: The short-term effect of the cash transfers is the CBR by age for the cross-section of women living at the time of the implementation of the policy, relative to the baseline (without cash transfers). The long-term effect is the CBR by age among women that had access to the policy in every period in which they could have children, relative to the baseline.

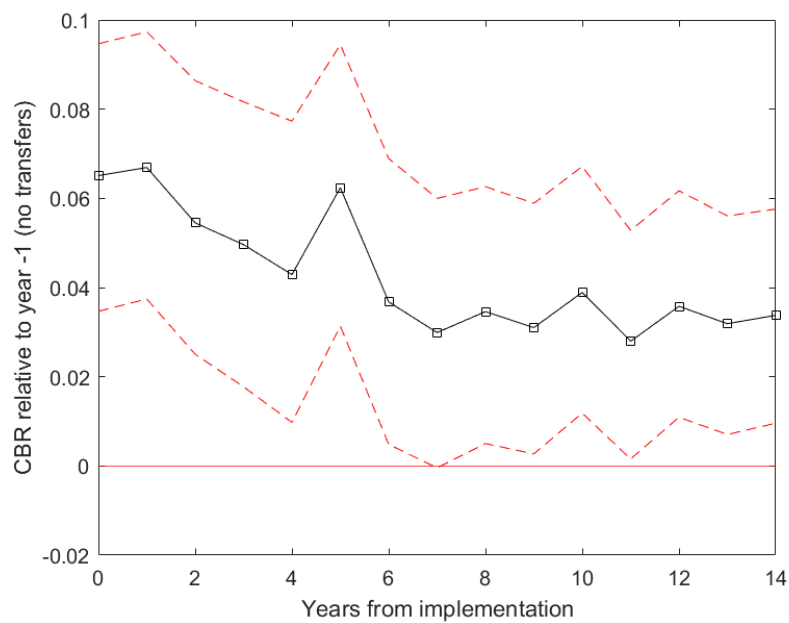
cash transfers are in place. The average age at first birth is reduced with the transfers by 0.15 years (about two months), while the average spacing between children is also shortened, but the magnitude of the change is smaller (less than a month between the first and second and slightly more between second and third).

In a nutshell, the main reason why the short and long-run impacts of the cash transfers are different boils down to the difference between the fertility response that women in different stages of the life-cycle have to the surprise announcement and the response when they have had time to adjust the timing of their fertility. For younger women, that response is not that different. For older ones, it is, because they cannot go back in time and change their previous decisions and have only a few periods to adjust. Over time, this additional effect from older women having to adjust immediately goes away, as they have had time to adjust.

5.2 Childcare subsidies

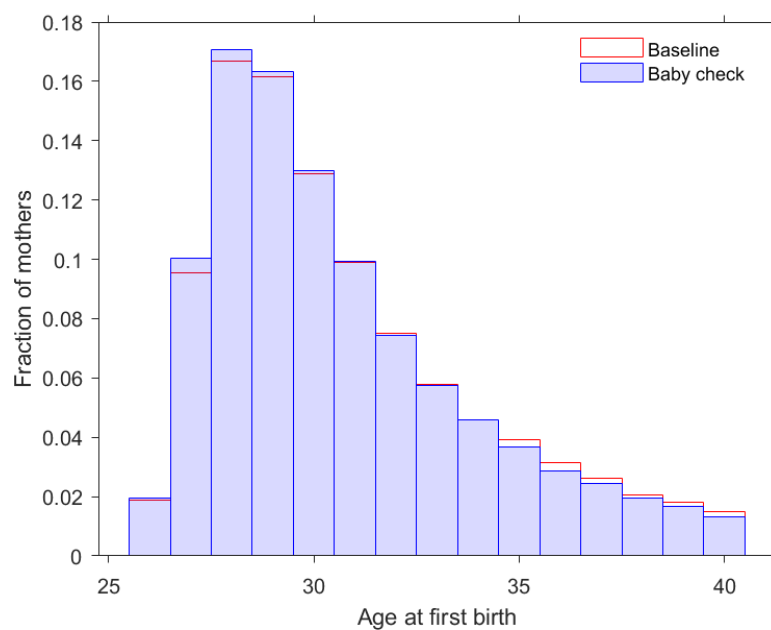
Childcare availability (or lack thereof) is frequently cited as one of the reasons why fertility may be low. Subsidising childcare is a natural alternative to a universal cash transfer, such

Figure 11: Overall crude birth rate relative to no cash transfers by year after implementation



Note: Dashed red lines are 95% confidence intervals, constructed via bootstrapping.

Figure 12: Age at first birth in the model, baseline and with baby check



as the baby check. Here, we consider the effects of a subsidy for children aged 0-3, since schooling after that age is provided for free in Spain.

To be comparable in magnitude to the payment offered by the baby checks, we compute the proportional subsidy with a present discounted value equal to the baby check payment, i.e. €2500, for a full-time working mother. Such proportional subsidy is 43.34% of the yearly cost of full-time childcare (€2000), for the first three years of an infant’s life. Since, in our model, women must buy childcare for the amount of time they work, an alternative way of seeing this policy is a cash transfer conditional on working.

Not surprisingly, the long-term fertility effects are smaller than the ones observed with the baby checks. Completed fertility increases by only 2.06%, compared with 3% with cash transfers. The two policies diverge in their effects on labour force participation: it drops by a little bit more than 1 percentage point with the cash transfers with respect to the baseline, but it increases by almost 1 percentage point with the childcare subsidies. This result, again, is hardly surprising since the latter is essentially a subsidy to working mothers. Moreover, part-time work remains unchanged, meaning that all of the effects operates through full-time.

Finally, in terms of the effect of the childcare subsidies on timing, again, they are a slightly more muted than with the cash transfers. The average age at first birth decreases by close to a month and a half, while the spacing between children remains unchanged.

5.3 Impact of labour market duality

Temporary contracts are associated with delayed and depressed fertility. While most of the literature exploring the effect of this type of working conditions on fertility centers on uncertainty and stability, we highlight another mechanism: returns to experience at critical ages. Most women start their careers under a temporary contract and anticipate that their labour market participation may be reduced after having their first child. Therefore, they have an additional incentive to work full-time, accumulate experience and postpone childbirth, i.e., obtain a permanent contract first.

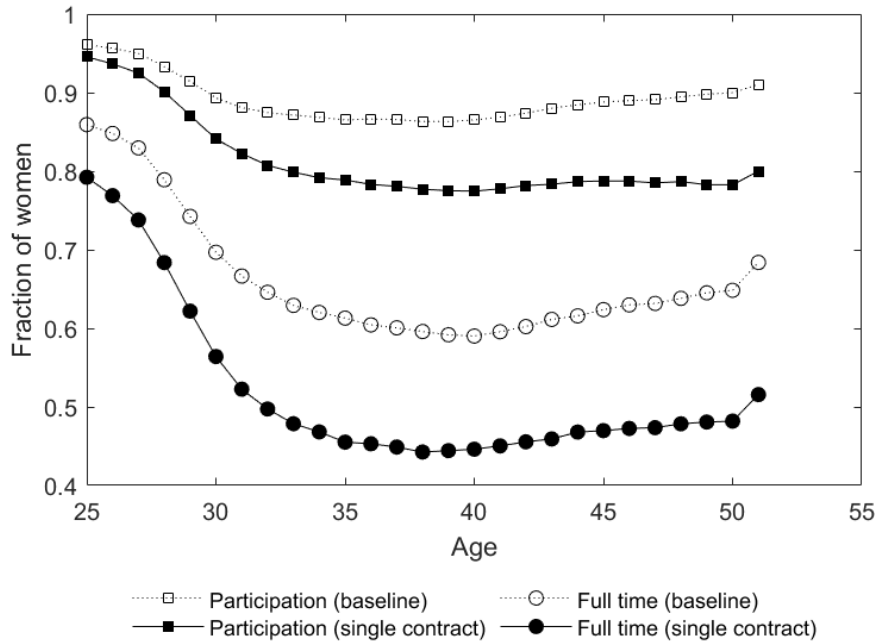
In our model, the difference between temporary and permanent contracts is reflected in the parameters of the income process. In particular, the constant term is larger for the permanent one, reflecting the fact that temporary workers experience unemployment spells during the year. Moreover, the returns to experience are larger for permanent workers, which could reflect firm or tenure-specific skill accumulation that temporary workers do not accrue.

The counterfactual we propose here is to have a single contract represented by a unique income process with parameters estimated from pooling all female workers. Among younger workers, this contract offers full-time earnings that are lower than the permanent contract, but higher than the temporary one in the baseline scenario. Therefore, the returns to experience are not as large at the beginning of women’s careers.

We re-compute the model solution under the new parametrisation for the income process, simulate the life-cycle for cohorts of women that live under these labour market conditions, and compare it to the baseline scenario (with dual labour markets and no cash transfers). Our main result is a 9.6% increase in completed fertility rates, which is three times as much as the long-term increase we found with cash transfers. Moreover, the first age at birth is anticipated by 5 months, and the timing between births is slightly shortened.

To understand better what is driving these effects, figure 13 shows the average overall and full-time labour force participation in the baseline and under the single contract. The former is not very different early in women’s working lives, but there is a gap in full-time participation early on. This gap widens as women get children (which happens earlier in the single contract scenario).

Figure 13: Female labour force participation through the life cycle, baseline and single contract



Finally, the last experiment we carry out is to implement the cash transfers on top of the single contract, and repeat the simulations. The completed fertility rate increases by 2.6% with respect to the scenario with a single contract but no transfers. This is very close but slightly lower than the effect the transfers had in the economy with dual labour markets. In terms of timing, the average age at first birth is further reduced by 0.12 years, or about a month and a half, with further smaller reductions in spacing, again very close to the effect on

timing on the economy with temporary and permanent contracts. These results imply that the results of our first experiment, i.e. that the effects on fertility of giving cash transfers upon birth are smaller in the long-run than in the short-run, and modest in magnitude, were not heavily dependent on that feature of the model and the Spanish economy.

6 Conclusions

A natural experiment involving cash transfers upon birth took place in Spain in 2007. The experiment was exploited by [González \(2013\)](#) via a DiD-RDD design to estimate its effects on fertility and mothers' labour force participation. The causal relationship identified thus is best interpreted as short-term, around the policy intervention.

In this paper, we develop a life-cycle model of fertility and labour force participation, and calibrate its parameters using these cleanly identified short-term effects, along with other aggregate moments for Spain in the period right before the implementation of the policy. Using the model, we explore the longer-term effects of the transfers, as well as the effects of alternative policy interventions, and their interactions with an important feature of the Spanish labour markets: the coexistence of temporary and permanent contracts (duality).

We find that the long-term effect of cash transfers on fertility is about half as large as it is in the short-term. The main reason is that in the short term there are additional births by older women that do not have time to adapt their previous fertility choices and have to adjust soon. Moreover, we find that a childcare subsidisation policy that gives women the same amount in present value as the cash transfers do brings about an increase in long-term fertility that is only slightly lower, but increases labour force participation instead of decreasing it. Furthermore, we find that the duality in the Spanish labour markets has a large effect on fertility, about three times as big as the effect of the transfers, driven by an increase in returns to experience during crucial years for child-rearing. However, labour market duality does not seem to be driving our results regarding the short versus long-term effects of cash transfers on fertility, as evidenced by the fact that the results change very little when we implement the policy in a scenario where only a single contract is available.

Our results highlight the importance of policy interventions that make labour market interruptions less costly (elimination of duality) and less necessary (childcare subsidies). This is in line with the main idea of the new economics of fertility, i.e. a crucial driver of it nowadays is the compatibility between career and family [Doepke et al. \(2022\)](#).

Appendices

A Computation of the model

A.1 Computation with taste shocks

All of the decisions that women make in the model are discrete in nature. Unfortunately, this type of models tend to generate jerky aggregate response to parameter changes, since individuals tend to change their decisions all at once, unless there are vast amounts of heterogeneity.

To facilitate the numerical solution of the model, we include a taste shock to women's utility in every period. This helps by smoothing out the labour force participation and fertility decisions. The shocks can be interpreted as unobserved state variables that add noise to the women's decisions. Moreover, the calibration and results are robust to their inclusion. For an in depth discussion of this computational method, see [Iskhakov et al. \(2017\)](#).

Thus, we assume that in every period women receive a vector of additive-separable taste shocks μ . In periods when they can still have children and need to choose on pregnancy $b \in \{0, 1\}$ in addition to labour force participation $h \in \{0, \frac{1}{4}, \frac{1}{2}\}$, they receive a vector of six shocks, one for every element in $\{0, \frac{1}{4}, \frac{1}{2}\} \times \{0, 1\}$. In periods when they cannot have any more children and need only to choose labour force participation, they receive a vector of three shocks, one for every element in $\{0, \frac{1}{4}, \frac{1}{2}\}$:

$$\mu = \begin{cases} \left(\mu_{0,0}, \mu_{\frac{1}{4},0}, \mu_{\frac{1}{2},0}, \mu_{0,1}, \mu_{\frac{1}{4},1}, \mu_{\frac{1}{2},1} \right) & \text{if } j < 15 \text{ and } N(\mathbf{n}) < 3 \\ \left(\mu_0, \mu_{\frac{1}{4}}, \mu_{\frac{1}{2}} \right) & \text{otherwise} \end{cases}$$

All of these shocks are i.i.d, drawn from an Extreme Value Type I distribution with scale parameter σ_μ .

The modified value function in states $(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)$ is:

$$W_j(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) = \begin{cases} \max\{W_j^{h,b}(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) + \sigma_\mu \mu_{h,b}\}_{h \in \{0, \frac{1}{4}, \frac{1}{2}\}, b \in \{0,1\}} & \text{if } j < 15 \text{ and } N(\mathbf{n}) < 3 \\ \max\{W_j^h(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) + \mu_h\}_{h \in \{0, \frac{1}{4}, \frac{1}{2}\}} & \text{otherwise,} \end{cases}$$

where $W_j^{h,b}$ and W_j^h represent the value, ex-taste shock, of choosing labour force participation h and pregnancy status b for a woman in period j , or just labour force participation h , in states $(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)$:

$$\begin{aligned} W_j^{h,b}(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) &= u^{h,b}(c, l, \mathbf{n}; N^*) + \beta \mathbb{E}^{\sigma_\mu} [W_{j+1}(\epsilon^{f'}, \epsilon^{m'}, z', x', \mathbf{n}', \mu; N^*)] \\ W_j(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) &= u^h(c, l, \mathbf{n}; N^*) + \beta \mathbb{E}^{\sigma_\mu} [W_{j+1}(\epsilon^{f'}, \epsilon^{m'}, z', x', \mathbf{n}', \mu; N^*)], \end{aligned}$$

where \mathbb{E}^{σ_μ} denotes the expectations over future taste shocks and in both cases, choice and states variables need to be retrieved from the constraints and laws of motion:

$$\begin{aligned} l &= 1 - h - \xi(\mathbf{n}) \\ c &= I - \lambda 2h(n_1 + n_2) \\ I &= y^m + 2hy^f \left(1 - \mathbb{1}_{\{h=\frac{1}{2}\}}\phi\right) - T(y^m, y^f, h) \\ \ln(y^f) &= \eta_0^f + \Delta\eta_0^f \mathbb{1}_{z=1} + \left(\eta_1^f + \Delta\eta_1^f \mathbb{1}_{\{z=1\}}\right)x + \left(\eta_2^f + \Delta\eta_2^f \mathbb{1}_{\{z=1\}}\right)x^2 + \epsilon^f \\ \ln(y^m) &= \eta_0^m + \eta_1^m(j-1) + \eta_2^m(j-1)^2 + \epsilon^m \\ \epsilon^{f'} &= \phi^f \epsilon^f + \nu^f \\ \epsilon^{m'} &= \phi^m \epsilon^m + \nu^m \\ \mathbf{n}' &= \Lambda_j(\mathbf{n}, b) \\ x' &= \Pi_x(x, h) \\ z' &= \Pi_z(x, h, z). \end{aligned}$$

The main consequence of introducing the taste shocks is that the policy function becomes probabilistic. Given the distribution assumed for them, the probability that a woman chooses pregnancy decision b and labour force participation h in states $(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)$ when $j < 15$ and $N(\mathbf{n}) < 3$ is the logit probability:

$$P_j(h, b \mid \epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) = \frac{\exp\left(\frac{W_j^{h,b}(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)}{\sigma_\mu}\right)}{\sum_{i \in \{0,1,2,3\}} \sum_{k \in \{0,1\}} \exp\left(\frac{W_j^{i,k}(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)}{\sigma_\mu}\right)}.$$

Otherwise, the probability that a woman chooses labour force participation h in states $(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)$ is the logit probability:

$$P_j(h \mid \epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*) = \frac{\exp\left(\frac{W_j^h(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)}{\sigma_\mu}\right)}{\sum_{i \in \{0,1,2,3\}} \exp\left(\frac{W_j^i(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)}{\sigma_\mu}\right)}.$$

One additional benefit of using Extreme Value Type I shocks is that the expected value function is given by the tractable log-sum formula from (McFadden, 1973):

$$\mathbb{E}^{\sigma_\mu} [W_{j+1}(\epsilon^{f'}, \epsilon^{m'}, z', x', \mathbf{n}', \mu; N^*)] = \begin{cases} \sigma_\mu \log \left(\sum_{i \in \{0,1,2,3\}} \sum_{k \in \{0,1\}} \exp \left(\frac{W_j^{i,k}(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)}{\sigma_\mu} \right) \right) & \text{if } j < 15 \text{ and } N(\mathbf{n}) < 3 \\ \sigma_\mu \log \left(\sum_{i \in \{0,1,2,3\}} \exp \left(\frac{W_j^i(\epsilon^f, \epsilon^m, z, x, \mathbf{n}; N^*)}{\sigma_\mu} \right) \right) & \text{otherwise.} \end{cases}$$

Using backward induction starting in period J , one can easily retrieve the expected value functions and the probabilistic policy functions.

B Stochastic structure for children's ageing

In this appendix we show how to retrieve the stochastic structure that governs the transition probabilities for the vector state of number of children at different ages takes, which we denote by $\mathbf{n}' = \Lambda_j(\mathbf{n}, b)$.

We denote by λ_1 and λ_2 the probabilities that an individual baby becomes a school-age child, and that an individual school-age child becomes a teen in a given period, respectively. Moreover, we assume that the ageing event is independent across children. Notice that $b\alpha_j$ is the probability that there is a newborn in the next period. Denote by $P_i(x \mid n_i)$ the probability that x children in stage $i \in \{1, 2\}$ (babies, school-age) move on to the next stage the next period (school-age, teenager), conditional on there being n_i children in that stage in the current period. Table 5 shows these, for babies and school-age children.

Table 5: Probabilities of ageing by number of children, $P_i(x \mid n_i)$

| n_i | Number of children ageing | | | |
|-------|---------------------------|------------------------------|------------------------------|---------------|
| | 0 | 1 | 2 | 3 |
| 0 | 1 | 0 | 0 | 0 |
| 1 | $1 - \lambda_i$ | λ_i | 0 | 0 |
| 2 | $(1 - \lambda_i)^2$ | $\lambda_i(1 - \lambda_i)$ | λ_i^2 | 0 |
| 3 | λ_i^3 | $\lambda_i(1 - \lambda_i)^2$ | $\lambda_i^2(1 - \lambda_i)$ | λ_i^3 |

To compute the whole set of probabilities of transition from state $\mathbf{n} = [n_0, n_1, n_2, n_3]$, we follow the algorithm:

```

for  $x_1 \in \{0, 1, 2, 3\}$  do
  for  $x_2 \in \{0, 1, 2, 3\}$  do
    if  $n_0 = 1$  then
       $\mathbf{n}' = [0, n_1 - x + 1, n_2 + x - y, n_3 + y]$  w.p.  $(1 - b\alpha_j)P_1(x_1 \mid n_1)P_2(x_2 \mid n_2)$ 
      or  $\mathbf{n}' = [1, n_1 - x + 1, n_2 + x - y, n_3 + y]$  w.p.  $b\alpha_jP_1(x_1 \mid n_1)P_2(x_2 \mid n_2)$ ;
    else
       $\mathbf{n}' = [0, n_1 - x, n_2 + x - y, n_3 + y]$  w.p.  $(1 - b\alpha_j)P_1(x_1 \mid n_1)P_2(x_2 \mid n_2)$  or
       $\mathbf{n}' = [1, n_1 - x, n_2 + x - y, n_3 + y]$  w.p.  $b\alpha_jP_1(x_1 \mid n_1)P_2(x_2 \mid n_2)$ ;
    end
  end
end

```

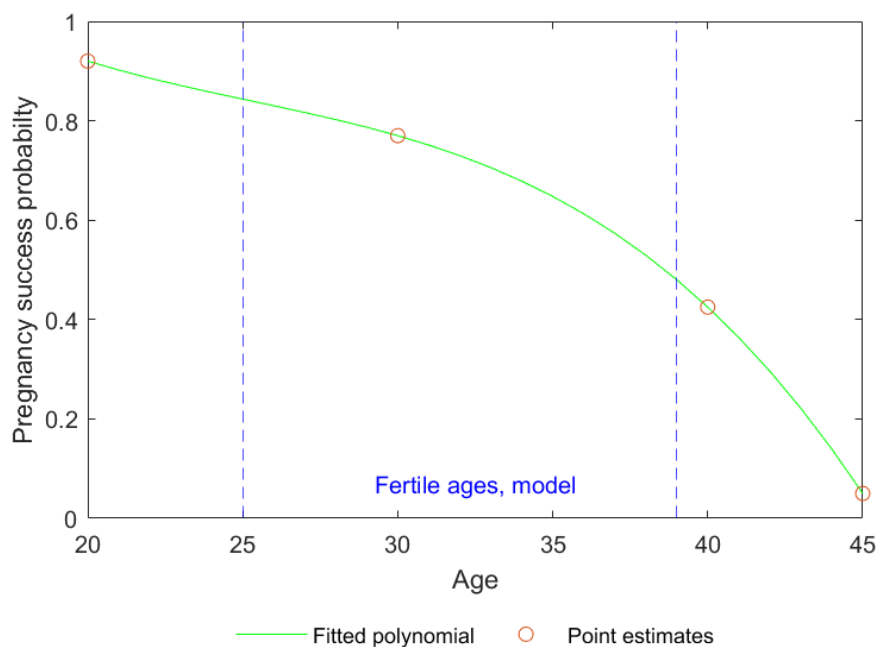
C Pregnancy success probabilities

We follow [Sommer \(2016\)](#) in estimating the probability of pregnancy success by age α_j . We use the following point estimates of natural infertility from [Trussell and Wilson \(1985\)](#):

$$\text{Infertility probability} = \begin{cases} 0.080 & \text{at age 20} \\ 0.230 & \text{at age 30} \\ 0.575 & \text{at age 40} \\ 0.950 & \text{at age 45,} \end{cases}$$

and then fit a polynomial through (the inverse of) them, as shown in figure 14. Notice that we only have women having children in the model between ages 25 and 39.

Figure 14: Pregnancy success probability conditional on age



Source: Author's work, point estimated by [Trussell and Wilson \(1985\)](#).

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