Maternity Leave and Female Labor Force Participation: Evidence from 159 Countries*

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Abstract

In this paper, we account for the direct and indirect effects of maternity leave entitlements on female labor force participation. We first explore theoretically the impact of maternity leave duration on female labor supply in the presence of fertility decisions. We assume that maternity leave duration affects female labor supply through two main channels: reducing the time cost of female market work, and reducing women's earnings. Our theoretical model allows for non-monotonic effects of leave duration on female labor supply. We test the predictions of our model using an unbalanced panel of 159 countries for the years 1994, 2004 and 2011. We confirm the existence of an inverted U-shaped relationship between maternity leave duration and female participation, and find a maternity leave threshold of around 30 weeks above which female participation falls. Below this threshold, increasing maternity leave increases female labor force participation because the positive effect due to the reduction of work-time cost of employed mothers strongly dominates the negative wage penalty effect. Beyond this threshold, the opposite happens. Our analysis also confirms the relevance of social norms for female labor supply throughout the world.

Keywords: female labor force participation, fertility, maternity leave, social norms.

JEL Classification: D13, J13, J16, J22.

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1 Introduction

It is generally acknowledged that the impact of maternity leave on labor market outcomes depends, in part, on the duration of leave. Using data for nine European countries over the period 1969-1993, Ruhm (1998) shows that a parental leave of around three months is associated with increases in the employment of women, especially those of child-bearing age, while lengthier leaves (around nine months) have no impact on employment but are associated with lower female hourly earnings. Thévenon and Solaz (2012) exploit a sample of 30 OECD countries between 1970 and 2010 and find that parental leave below two years has a positive effect on female employment, while longer leaves reduce employment. For the U.S., Rossin-Slater (2017) reports non-monotonic effects of family leave policies on women's employment rates: while leave entitlements less than one year in length improve their employment rates, longer leaves can have negative effects. Focusing on Canada, Baker and Milligan (2008) find that leave entitlements of 17-18 weeks don't change the amount of time women spend away from work, but longer leaves of up to 70 weeks lead them to spend more time at home. Olivetti and Petrongolo (2017) review this extensive literature and add to cross-country evidence with data from 30 OECD countries from 1970 to 2014. They confirm the existence of a non-monotonic effect such that the employment rate rises with leaves up to 50 weeks and declines thereafter.[1]

Job-protected maternity leave can also impact fertility. Olivetti and Petrongolo (2017) report that the effect of parental leave on fertility is non-monotonic (increasing more than proportionally as the length of the leave increases), but quantitatively negligible in OECD countries. Still, Averett and Whittington (2001) find positive effects of parental leave on fertility in the U.S., as does Björklund (2006) in Sweden. Lalive and Zweimüller (2009) find strong positive effects on fertility due to an extension of paid protected leave in Austria in the 90s, and Raute (2019) also reports evidence from Germany suggesting that leave policies can affect fertility. Importantly, leave policies can also have an indirect effect on female labor force via fertility. In relation to this, Aaronson et al. (2018) and Bloom et al. (2009) report that fertility may lower female labor participation rates.

To account for both the direct and indirect effects of maternity leave duration on female participa-

[1]Studies have also considered the effect of leave duration on outcomes in the short and long-term. In general, longer leaves reduce female employment in the time immediately after birth as recent mothers take advantage of the eligibility, but have a limited or negligible impact on female participation rates in the long run (see, for example, Hanratty and Trzcinski (2009) for the case of Canada, Lalive and Zweimüller (2009) for Austria, Schönberg and Ludsteck (2014) for Germany and Dahl et al. (2016) for Norway). We do not consider the short- versus long-run effect of leave duration in this article. Our results should be interpreted as average effects across time and households.

tion rates, we extend previous work in two ways. We first explore theoretically the impact of maternity leave duration on female labor supply in the presence of fertility decisions. We take as our starting point Bloom et al.'s (2009) unitary model, where households choose both women's labor supply and the number of children. We extend this model by including maternity leave entitlements that reduce the time cost of female market work, may reduce the wage received during their absence, and are costly to the firm.[2] Overall, maternity leave affects decisions through two main channels: reducing the time cost of female market work, and reducing women's earnings. The first channel has a positive direct effect on female labor supply but, indirectly, also reduces labor supply because a lower time cost of female market work raises fertility. The second channel reduces labor supply when the substitution effect of lower wages dominates. Thus, our theoretical model allows for non-monotonic effects of leave duration on female labor supply, including its indirect effect through fertility decisions. To the best of our knowledge, the existing theoretical literature has not explored the relationship between maternity leave duration and labor supply in the presence of fertility decisions.

Our theoretical model also extends Bloom et al. (2009) by incorporating the role of social norms in women's labor supply and fertility decisions. We consider social norms concerning the extent to which it is appropriate for women to work, and assume that they affect each household's utility in two ways. First, there is a specific cost for women of finding a job that decreases as views towards working women are more favorable. Second, the social norm affects the costs of avoiding fertility, as suggested in Goldin and Katz (2002). Thus, we assume that where views towards working women are more favorable, the cost of avoiding fertility is lower. Interestingly, our model shows that, once fertility decisions are internalized, social norms have an ambiguous effect in labor supply. Specifically, if social views towards working women become more favorable, there is a direct positive effect on female labor supply because the time cost of finding a job decreases. However, an ambiguous indirect effect via fertility appears: on the one hand, lower search costs can increase fertility and thus decrease female labor supply; on the other hand, the cost of reducing fertility decreases, resulting in fewer children

[2] Job absence due to maternity leave can generate direct costs for firms in the guise of replacement and overtime costs, and indirect ones in the form of productivity losses. Survey information provided by SHRM and Kronos (2014) shows that the direct costs and the productivity loss of a planned absence related to vacations, sick time and parental leave are equal to 29.4% and 15.2% of total payroll in Europe, respectively (see Tables 11 and 12 in the report). The report also calculates similar costs in the U.S., Australia, China, India and Mexico. Additionally, as mentioned by Thévenon and Solaz (2012), if employees take up long leave entitlements, they may become detached from the labor market as their skills depreciate. For these reasons, we assume that the wage penalty associated with the productivity loss depends on the length of the leave awarded to mothers and that its effect is quadratic.

and thereby increasing female labor supply.

Our second contribution is empirical. Specifically, unlike previous cross-country evidence that has focused on samples of European or OECD countries, we test for the existence of a non-linear relationship between maternity leave duration and female labor market participation rates in a wide sample of 159 countries. We test the theoretical predictions of our model using an unbalanced panel covering the years 1994, 2004 and 2011. We estimate a set of ordinary least square (OLS) regressions and, consistent with our theoretical priors, we report the existence of an inverted U-shaped relationship between female labor supply and maternity leave across countries. Specifically, we find a maternity leave threshold of around 30 weeks above which female participation falls. According to the theoretical model, to the left of this threshold, increasing maternity leave increases female labor force participation because the positive effect due to the reduction of work-time cost of employed mothers strongly dominates the negative wage penalty effect. Beyond 30 weeks, the wage penalty effect is stronger, thus reducing female labor supply.

Finally, the regressions confirm the relevance of social norms for female labor force participation (e.g.: Akerlof and Kranton; 2010; Bertrand, 2011; Fernandez, 2013; Bertrand et al., 2015; Bertrand et al., 2016; Codazzi et al., 2018; Petrongolo, 2019; Casarico and Profeta, 2020). In the words of Bertrand (2011, page 1571): "as long as there is a strong behavioral prescription indicating that men work in the labor force and women work in the home, norms regarding gender identity could explain why women have been slow at increasing their labor force participation". Our estimates indicate that an increase of 10 percentage points (about one standard deviation) in the prevalence of positive attitudes towards working women increases the female labor force participation by almost 6.8 percentage points. As suggested by our theoretical model, the reduction in both the time cost of finding a job and the cost of reducing fertility may be behind this positive effect.

The rest of the paper is organized as follows. In Section 2 we review the theoretical work that has linked fertility and labor supply decisions. We then present the theoretical model linking maternity leave to female labor supply in Section 3. In Section 4 we explain how we measure our key variables, and in Section 5 we present and discuss our main empirical results as well as related robustness analysis. We conclude the article with Section 6.

2 Theoretical literature

Female labor market outcomes and fertility are clearly interconnected. The modeling of women's labor supply and fertility as conflicting simultaneous decisions has a long history in the theoretical literature. Following the tradition of Becker's unitary model, Cigno (1986) considers households as single decision units characterized by a utility function of joint adult consumption, number and quality of children. This generates important insights into the design of family policies. It points, for instance, to a potential trade-off between the number and quality of children in the absence of externalities. Cigno (1998) extends this work to allow for fertility adjustments accounting for infant mortality.

Galor and Weil (1996) consider an overlapping generations general equilibrium model where men supply inelastically one unit of mental labor and one unit of physical labor and women, endowed with less physical labor, choose how much mental labor to supply. There is no leisure: households decide only how many children to have and labor supply adjusts automatically. This paper predicts a negative relationship between income, positively associated with women's labor force participation, and fertility. However, in poor countries, households are constrained in the number of live births. When income increases in these countries, more children survive and women spend more time childrearing and less in labor-force participation. This replicates the U shape relationship between female labor supply and income in Goldin (1994): when output is low, an increase in income is related to less labor-force participation (and more children), when output is high more income is related to more labor force participation (and less children).

Apps and Rees (2004) note that the traditional negative relationship between female labor supply and fertility becomes positive in the 90s. They propose a very simple model where men supply one unit of labor inelastically and women allocate their time to market work and childcare, and use it to attribute the change in tendency to a change to public policy from joint to individual taxation, and from child-related (cash) transfers to child-care (in-kind) transfers. Bloom et al. (2009) extend this simple model to account for women's leisure, child mortality, fertility control and the role of urbanization, and identify the (positive) causal effect of fertility reductions on female labor force participation.

In spite of these achievements, the unitary model has clear limitations. For example, because all incomes are pooled together, it fails to account for empirical regularities concerning the adjustment of both fertility and labor supplies to changes in incomes allocated to one or the other member of the household (e.g. Schultz, 1990). This limitations are still more patent when it comes to conducting normative welfare analysis, leaving no room to determine the optimal allocation of consumption, labor supply, housework, in sum, welfare, between the members of the household. Thus, another strand of the literature has diverged from the unitary model by recognizing that individual preferences within the household may differ.

Decision-making within this (game theoretical) framework can be assumed to follow a non-cooperative (Cigno, 2012) or a cooperative (Chiappori, 1997) pattern. In the former case, each member of the household maximizes her utility given the choices of the other member. Because this can lead to Pareto inefficient intra-household allocations, the use of cooperative models based on bargaining theory (the collective approach) has received more support (see Vermeulen (2002) and Chiappori et al. (2019)). In these models, members of the household have distinctive preferences and cooperate in, or bargain over the production of a common household good or public good. Although the public good can certainly be interpreted as children like in Blundell et al. (2005), fertility decisions are generally not explicit in these models. Eswaran (2002) considers fertility but not labor supply decisions.

As we have mentioned before, our model goes back to the unitary tradition and extends Bloom et al. (2009) in several directions. Despite the aforementioned limitations of this unitary model, our predictions concerning the relationship between maternity leave duration and labor supply are strongly supported by the empirical evidence. We present the model in detail next.

3 The Model

The model is based on Bloom et. al (2009). A household consists of a man, who supplies inelastically one unit of efficient labor, and a woman that divides her time among work, childcare and leisure.

The utility function of a representative household is defined over household consumption c, female leisure d, and fertility n, and is given by

$$U(c,d,n) = \log{(c - c_0)} + \alpha \log{(d)} + \eta(1-m)n - k\left(\Sigma, \Lambda\right)(N-n) \tag{1}$$

For simplicity it is assumed that utility is logarithmic in consumption and leisure and linear in the number of surviving children $(1-m)n$ where m is the infant mortality rate. The weight on consumption in utility is normalized to 1, the relative weight of leisure is $\alpha > 0$, and the relative weight given to surviving children is $\eta > 0$. Parameter c_0 represents subsistence consumption. The potential reproductive capacity (usually 15 pregnancies) can be reduced at a cost k that represents the cost of actions such as abstinence, using contraceptives, or interrupting pregnancies. While Bloom et al. (2009) take this cost as given, we allow it to be a function of the social norm parameter Σ and other, potentially exogenous, factors Λ, $k(\Sigma, \Lambda)$.[3] Parameter Σ measures the general attitude towards working women, with 0 meaning least and 1 most favorable towards working women. Our assumption is that the more favorable are views towards working women in a country the less costly it is to avoid fertility therein ($k_\Sigma < 0$). Parameter Λ can collect diverse elements with positive or negative effects on the costs of reducing fertility but as in Bloom et al. (2009) we will consider the extent of legal grounds permitting abortion that will have a negative effect on the costs of reducing fertility ($k_\Lambda < 0$).[4]

Total time available to a woman during the year is normalized to 1. This is divided between working time l_f, leisure time d, and childcare. Time allocated to childcare is linear in the number of surviving children $n(1-m)$ where n stands for fertility and m for infant mortality rates, with a time cost per child of b. There is a work-time-cost per unit of labor supplied, ϕu, where $\phi > 0$ and u is an indicator of urbanization. This time cost is meant to account for commuting time, which is larger the larger the size of the city. Having to commute for a longer time implies more time devoted to supplying the same amount of labor.

Going beyond Bloom et al. (2009), we assume that urbanization has a positive effect on income and wages, as will be made clear below. In addition, we introduce a time cost of finding a job ψ that is normalized to zero in the case of men and depends negatively on the social norm parameter Σ in the case of women. Thus, the specific cost of finding a job for a woman $\psi(\Sigma)$ decreases as views towards working women are more favorable: $\psi'(\Sigma) < 0$. Finally, we model parental maternity leave mandates $\lambda < 1$ as reducing the work-time-cost of employed mothers.[5] In other words, a maternity leave of duration λ represents λl_f more time available for mothers on a contract for l_f hours. Summing up, the yearly leisure constraint of a typical woman is:

$$d = 1 - (1 - \lambda + \phi u + \psi(\Sigma))l_f - bn(1-m) \tag{2}$$

[3] Goldin and Katz (2002) point to the role of social norms in maintaining laws that prohibited the sale of contraceptives in the U.S. in the middle of the 20th century. Other things, such as country imitation effects or pressure from international organizations, could have an independent effect on the costs of reducing fertility k.

[4] Although it can be argued that this too affects social norms, we will allow for Λ to affect household decisions through an additional independent channel.

[5] Parental leave rights can only be used once a year and because of this, we do not multiply this term by the number of surviving children.

As in Bloom et al. (2009), men supply inelastically one unit of labor and receive the male wage w_m in exchange. Women supply l_f hours of work and earn the wage w_f while not on leave. We assume that, while on leave, they earn a proportion ρ of their wage. Household consumption possibilities are therefore given by

$$c = w_m + w_f l_f \left(1 - \lambda \left(1 - \rho\right)\right) \tag{3}$$

Substituting (2) and (3) into (1) we obtain the indirect utility function:

$$\begin{aligned} V(l_f) &= \log\left[\left(w_f l_f \left(1-\lambda\left(1-\rho\right)\right) + w_m\right) - c_0\right] + \alpha \log\left[1 - (1 - \lambda + \phi u + \psi(\Sigma)) l_f - bn(1-m)\right] \\ &\quad + \eta(1-m)n - k(\Sigma, \Lambda)(N-n) \end{aligned} \tag{4}$$

We first assume that fertility decisions are exogenous and study the choice of female labor supply for n given. Then, in section 3.1 we account for the endogeneity of fertility decisions.

To estimate wages from human capital achievement (years of schooling), we use a Cobb-Douglas production function:

$$Y = A(u) K^{\sigma} E^{1-\sigma}$$

where $A(u)$ is an increasing function of the urbanization indicator u, accounting for positive agglomeration effects, K stands for capital, and E stands for efficiency units of labor:[6]

$$E = E_m h_m + E_f h_f (1 - \theta \lambda^2)$$

The term $\theta\lambda^2$ captures administrative costs and potential losses of productivity experienced by the firm when a worker is on temporary leave. Accounting for previous evidence mentioned in the Introduction, this effect is assumed to be convex on the duration of the leave, with $\theta > 0$. Firms are competitive and wages equal their marginal products:

$$w_m = (1-\sigma) y h_m \tag{5}$$

$$w_f = (1-\sigma) y h_f \left(1 - \theta\lambda^2\right) \tag{6}$$

with $y = Y/E = A(u)(K/E)^{\sigma}$.

We now derive the equations to estimate and provide their linearized form together with a discussion of the predicted signs of the coefficients (see the Appendix for further details).

[6]Congestion of local public goods can also arise when urbanization increases, resulting in a negative effect on A (Combes and Gobillon, 2015).

3.1 Female labor supply with exogenous fertility

From (4), and using (5)-(6), we obtain the first order condition determining female labor supply:

$$l_f^* = \frac{(1 - bn(1 - m))}{(1 + \alpha)(1 - \lambda + \phi u + \psi(\Sigma))} - \frac{\alpha((1 - \sigma)yh_m - c_0)}{(1 + \alpha)(1 - \lambda(1 - \rho))(1 - \sigma)yh_f(1 - \theta\lambda^2)} \tag{7}$$

where $y = A(u)(K/E)^\sigma$. Thus, in addition to the negative effect of urbanization on female labor supply through the reduction in available time in Bloom et al. (2009), there is a positive effect on wages due to agglomeration effects.

Linearizing (7) we can write:

$$l_f = \beta + \beta_y y + \beta_{hm} h_m + \beta_{hf} h_f + \beta_\rho \rho + \beta_\lambda \lambda + \beta_{\lambda^2} \lambda^2 + \beta_n n + \beta_m m + \beta_\Sigma \Sigma + \beta_u u \tag{8}$$

and we can show that: $\beta_y < 0$, $\beta_{h_m} < 0$, $\beta_{h_f} > 0$, $\beta_\rho > 0$, $\beta_\lambda \lesseqgtr 0$, $\beta_{\lambda^2} < 0$, $\beta_n < 0$, $\beta_m > 0$, $\beta_\Sigma > 0$, $\beta_u < 0$ (see Appendix). In words, income y, and the human capital of men h_m, both have a negative effect on female labor supply through an income effect. In contrast, the human capital of women h_f, and the wage replacement rate while on leave ρ, can increase female labor supply if the substitution effect of the higher wage dominates.[7] The linear effect of leave duration on the labor supply of women β_λ could be positive or negative when the number of children is given. This is due to the fact that the leave endows women with more time and this has a positive effect on labor supply, but may reduce earnings when working if $\rho < 1$, and this has a negative effect on labor supply. If $\rho = 1$, the duration of the leave will have a positive linear effect on female labor supply.[8] The negative effect of the square of leave duration λ^2 suggests and inverted U shape relationship between leave duration and female labor force participation. More children n (or lower child mortality m) reduce available time, and the social norm that reduces the cost of finding a job for a woman Σ, has a positive effect on the female labor supply when fertility is given. Finally, living in a larger city (represented by the urbanization parameter u) increases commuting costs, which reduces female labor supply, and generates higher income through agglomeration effects, which also reduce female labor supply. The reason for this is that the male wage increases, with an income effect that reduces female labor supply, and although the female wage also increases, with an income and a substitution effect, income effects can be shown to dominate (see Appendix). In the presence of congestion costs due to urbanization ($A'(u) < 0$, see

[7] As in Bloom et al. (2009) the substitution effect will dominate provided that $w_m > c_0$.

[8] The sign of the *marginal* effect of leave duration on female labor supply will however not only depend on β_λ but also on β_{λ^2} being $\beta_\lambda + 2\beta_{\lambda^2}$.

footnote 6), reduced income would imply a positive effect of u on female labor supply. Then the sign of β_u would be ambiguous.

We now consider the endogenous choice of fertility n.

3.2 Female labor supply with endogenous fertility

From (4), the optimal fertility choice will be given by:

$$n^* = \frac{(1-(1-\lambda+\phi u+\psi(\Sigma))l_f)}{b(1-m)} - \frac{\alpha}{(\eta(1-m)+k(\Sigma,\Lambda))} \tag{9}$$

Since fertility n depends on labor supply, and to account for this endogeneity, we first linearize (9) and then substitute the outcome in (8). The linear version of (9) is:

$$n = \gamma + \gamma_\lambda \lambda + \gamma_{lf} l_f + \gamma_m m + \gamma_u u + \gamma_\Sigma \Sigma + \gamma_\Lambda \Lambda \tag{10}$$

where $\gamma_\lambda > 0$, $\gamma_{lf} < 0, \gamma_m \lesseqgtr 0, \gamma_u < 0, \gamma_\Sigma \gtreqless 0$, and $\gamma_\Lambda < 0$ (see Appendix). A larger duration of the maternity leave has a positive effect on fertility as expected. Also as expected, a higher labor supply reduces fertility. Higher child mortality has an ambiguous effect (it increases available time but requires higher fertility to attain the desired number of children). Urbanization u increases the time cost of finding a job, reducing available time and hence fertility of working mothers and having no effect on women who do not work. More favorable views towards working mothers Σ decrease the time cost of finding a job for those women who want to work, thus increasing available time and, hence, fertility. However, such favorable views also decrease the cost of reducing fertility, and this has a negative effect on n. Thus the final effect of Σ on fertility is ambiguous in sign, but negative for non-working women. Finally, more permissive abortion laws Λ make abortion easier and hence reduce fertility.

We now substitute (10) into (8) and, collecting terms, obtain the equation that we estimate in the next section:

$$l_f = \mu + \mu_y y + \mu_{h_m} h_m + \mu_{h_f} h_f + \mu_\rho \rho + \mu_\lambda \lambda + \mu_{\lambda^2} \lambda^2 + \mu_m m + \mu_\Sigma \Sigma + \mu_\Lambda \Lambda + \mu_u u \tag{11}$$

The signs of these parameters are derived in the Appendix. We obtain: $\mu_y < 0$, $\mu_{h_m} < 0$, $\mu_{h_f} > 0$, $\mu_\rho > 0$, $\mu_\lambda \gtreqless 0$, $\mu_{\lambda^2} < 0$, $\mu_m \gtreqless 0$, $\mu_u < 0$, $\mu_\Sigma \gtreqless 0$, and $\mu_\Lambda > 0$. In the model, income y, the human capital of males and females h_m and h_f, the term λ^2, and ρ, do not affect the optimal choice of fertility. For this reason, the sign of the effect of these variables on female labor supply is the same as the one obtained when fertility is assumed exogenous. The linear effect of the duration of the leave λ on female labor supply, μ_λ, continues to be ambiguous but, now, it is so even if $\rho = 1$: a higher duration reduces the time cost of working, may reduce earnings (or not, if $\rho = 1$) and, now, also induces higher fertility which has a negative effect on female labor supply. Infant mortality m which had an unambiguous positive effect on labor supply when fertility was exogenous, now has an indeterminate influence because of the ambiguous effect of child mortality on fertility. Urbanization u has a direct negative effect on female labor supply and also affects the fertility choice negatively. For this reason, the final effect of u on female labor supply may be ambiguous. Still, the calculations displayed in the Appendix show that the global effect is still negative when we account for endogenous fertility.

Finally, the effect of the social norm Σ, clearly positive when fertility is taken as exogenous, is now ambiguous: $\mu_\Sigma \gtreqless 0$. If social views towards working women become more favorable, there is a direct positive effect on female labor supply because the time cost of finding a job decreases. But positive views towards female employment also have an ambiguous effect on female labor supply via fertility because, on the one hand, working women have more time to have children and, on the other hand, the cost of reducing fertility decreases resulting in less children. In contrast, the permissiveness of abortion laws Λ has an unambiguous positive effect on female labor supply ($\mu_\Lambda > 0$).

The rest of the paper will empirically test these theoretical predictions.

4 Data

As already stated in the Introduction, our data takes the form of an unbalanced panel covering the periods 1994, 2004 and 2011 and 159 countries. Table 1 provides summary statistics for all the variables employed in the article, including the overall-O, between-B (variation between countries) and within-W (variation within countries) standard deviations (SD). The dependent variable is the female labor force participation rate l_f taken from the International Labor Organization (ILO). The female labor force participation rate is the proportion of women aged between 15 and 64 years old who are economically active. It ranges from 8.8% to 89.8% with an average of 50.8%. Countries on the lower end of the distribution include Iran and Algeria while those with the highest rates are Iceland and Mozambique, among others. Most of the variation in this variable is between countries as attested by the overall SD of 17.5, a between SD of 17.3 and a within variation of 3.2.

We are interested in the effect of maternity leave, λ on female labor supply. We measure leave length by using the total duration (in weeks) of maternity leave whether paid or unpaid for singleton and uncomplicated childbirths taken from the ILO Working Conditions Laws Database. Leave duration λ has an average value of 14.1 weeks and ranges from 4 in Tunisia in 1994 and 2004, to 58.6 weeks in Croatia in 2004 and 2011. As was the case for female force participation rates, within country variation in the maternity leave indicator is rather limited: within countries SD 2.1 compared to between countries SD of 7.3 in the context of an overall SD of 6.8.

The explanatory variables in equation (8) are: the fertility rate n (births per woman) taken from the World Population Policies Database; social norms Σ based on the percentage of the population agreeing with the assessment that *It is perfectly acceptable for any woman in your family to have a paid job outside the home if she wants one* taken from Gallup and ILO (2017); the replacement maternity leave rate ρ refers to the percentage of previous earnings replaced by the maternity benefit over the length of the paid leave entitlement taken from the ILO Working Conditions Laws Database; the infant mortality rate m per 1000 live births taken from the World Bank database; the male and female years of schooling (h_m and h_f) from the Human Development Report; and the GDP per capita y based on purchasing power parity (PPP) in constant 2011 international dollars and applying logs.

When estimating Eq. (11), we replace fertility with a proxy of the cost of avoiding childbirth. Following Bloom et al. (2009), we include the abortion index from the World Population Policies Database, which measures the total number of legal grounds on which abortion is permitted (from 0 to 7). This index reflects legal provisions under which the Government permits induced abortion in the country and includes seven legal reasons for an abortion: to save the life of the woman; to preserve her physical health; to preserve her mental health; consequent on rape or incest; fetal impairment; economic or social reasons; and available on request. A value index of 0(7) means than none(all) of these laws apply to the country. The higher the number of legal grounds permitting abortion Λ, the lower the expected cost of reducing fertility k. This will reduce the number of children and, therefore, increase female labor force participation.

As explained more fully below, for robustness we also employ a range of additional control variables namely, general government consumption as a share of GDP from the World Bank, legal traditions (La Porta el al. 2008) and religious affiliations (North et al. 2013).

Table 1: Model variables: Summary statistics

Variable	N	Mean	O(SD)	B(SD)	W(SD)	Min	Max
Female labor force participation, l_f (%)	477	50.8	17.5	17.3	3.2	8.8	89.8
Maternity leave, λ (weeks)	429	14.1	6.8	7.3	2.1	4	58.6
Positive attitudes towards working women , Σ (%)	378	85.25	12.13	12.13	12.13	35	100
Replacement maternity leave rate, ρ (%)	419	87.1	21.3	20.3	7.8	0	100
Fertility rate (births per woman), n	476	3.3	1.7	1.7	0.5	1.2	7.8
Abortion index, Λ	466	3.9	2.5	2.4	0.6	0	7
Infant mortality rate, m (per 1000 live births)	477	36.0	33.3	31.3	11.5	2	168
Urban population rate, u (%)	477	54.9	23.2	23.0	3.4	7.0	100
Years of schooling male, h_m	431	7.8	3.0	2.9	0.8	1.1	13.9
Years of schooling female, h_f	432	7.0	3.5	3.5	0.9	0.2	13.8
GDP per capita, y (constant dollars 2011 PPS)	457	15,520	18,948	19,292	3,687	477	124,024
General government consumption (% of GDP)	424	15.8	6.0	5.4	2.8	1.8	36.6
Legal origins French code	471	0.456	0.499	0.499	0	0	1
Legal origins German code	471	0.032	0.176	0.176	0	0	1
Legal origins Scandinavian code	471	0.032	0.176	0.176	0	0	1
Legal origins socialist communist laws	471	0.178	0.383	0.383	0	0	1
Religious affiliation Protestants	465	0.185	0.226	0.226	0	0	0.93
Religious affiliation Catholic	468	0.297	0.383	0.383	0	0	0.94
Religious affiliation Eastern Orthodox	468	0.052	0.162	0.162	0	0	0.91
Religious affiliation Muslim	468	0.231	0.346	0.346	0	0	0.99

Note: Variables as defined in the main text. The database is available from the authors on request.

5 Empirical Results

In Table 2, we report the OLS estimates when regressing female labor supply on maternity leave and the variables included in the theoretical model. In column 1 we regress female participation on maternity leave without considering any additional variables. Columns 2 and 3 correspond to Equation (8) of the theoretical section 3 that assumes exogenous fertility. Column 3 augments the regression in column 2 by including a squared of (log) GDP per capita term to account for the possibility of a quadratic relationship between GDP per capita and female participation (Goldin 1994; Mammen and Paxson 2000; Bloom et al. 2009). Controlling for fertility allows us to consider the direct effect of maternity leave on female participation or, in other words, allows us to separate out the indirect effect of leave on participation passing through fertility. In column 4 we replace fertility with the abortion index, which Bloom et al. (2009) employ as an exogenous determinant of fertility rates. This corresponds to Equation (11) of the theoretical section, capturing the total effect of maternity leave on female participation rates, including its effect via fertility.

Consistent with our theoretical priors, the results clearly indicate that maternity leave λ is asso-

Table 2: Determinants of female participation rates estimation

Regressors:	OLS(1)	OLS(2)	OLS(3)	OLS(4)
Maternity leave (λ)	1.908***	0.548*	0.634**	0.637**
	(0.378)	(0.341)	(0.313)	(0.296)
Maternity leave squared(λ^2)	-0.028***	-0.011**	-0.011**	-0.011**
	(0.006)	(0.005)	(0.005)	(0.004)
Maternity leave threshold (weeks)	34.3	24.9	28.8	29.0
Fertility rate (n)	-	0.695	-0.833	-
	-	(1.065)	(1.067)	-
Abortion index (Λ)	-	-	-	1.629***
	-	-	-	(0.289)
Attitudes working women (Σ)	-	0.907***	0.814***	0.767***
	-	(0.068)	(0.068)	(0.065)
Maternity replacement rate (ρ)	-	0.017	0.020	0.025
	-	(0.030)	(0.028)	(0.028)
Infant mortality rate (m)	-	0.182***	0.121*	0.122**
	-	(0.064)	(0.062)	(0.051)
Urban population rate (u)	-	-0.367***	-0.285***	-0.193***
	-	(0.061)	(0.062)	(0.059)
Years of schooling male (h_m)	-	0.121	-0.411	-1.681**
	-	(0.929)	(0.866)	(0.832)
Years of schooling female (h_f)	-	3.477***	3.615***	4.146***
	-	(0.889)	(0.829)	(0.806)
log GDP per capita ($log(y)$)	-	-3.033**	-61.560***	-56.871***
	-	(1.412)	(9.853)	(8.756)
log GDP per capita squared ($log(y)^2$)	-	-	3.093***	2.813***
	-	-	(0.520)	(0.455)
Number of observations	429	330	330	322
Adjusted R-squared	0.058	0.587	0.640	0.672

Note: *, **, *** measures statistical significance at the 10, 5 and 1 percent levels respectively. All regressions include a constant and report robust errors.

ciated with female participation rates but that the association is not linear. In particular, we confirm the existence of an inverted U-shaped relationship and, more specifically, we find a maternity leave threshold of between 25 and 34 weeks above which female participation falls. According to the theoretical model, if one is on the left of this threshold, increasing maternity leave will increase female labor force participation because the positive effect due to the reduction of work-time cost of employed mothers strongly dominates the negative wage penalty effect. Alternatively, beyond this threshold, the latter effect will overwhelm that due to the reduction in the work-time cost.

In addition, we find that labor supply is positively associated mortality rate m. In turn, participation rates are higher in societies that are favorable to female participation in the labor market. This is in line with the theoretical expectation that social or cultural norms can reduce the search cost experienced by women as well as the cost of avoiding fertility. Thus, our estimates indicate that an increase of 10 percentage points (near to one standard deviation) in the prevalence of positive attitudes towards working women Σ, increases labor force participation by around 8 percentage points.

We also find a positive effect of female education h_f on labor supply due to its positive effects on female wages (thus implying a dominant substitution effect according to the model). Specifically, an additional year of schooling increases labor force by more than 4 percentage points (see column 4). Additionally, urbanization reduces female labor supply, suggesting the presence of a work-time-cost effect that increases with the level of urbanization. Finally, columns 3 and 4 confirm the quadratic relationship between economic development and female participation in the labor market.

As previously stated, the results reported in column 4, show the total effect of maternity leave on female labor participation rates, including its effect via fertility (Equation (11)). They suggest that incorporating the indirect effect on female labor supply going through fertility, has a negligible effect on the maternity leave threshold (28.8 in OLS(3) to 29.0 in OLS(4)), indicating that the positive direct effect of maternity leave duration on female participation dominates the negative indirect effect running through fertility rates.

6 Robustness Analysis

The limited within variation in female participation rates and maternity leave duration, leads us to eschew country fixed-effects in the empirical analysis. However, because we are aware that omitted variable bias may affect our estimates, we pursue the robustness of our results in several directions. First, we account for the impact of time varying variables common to our cross-section units through the application of period fixed effects, as well as the effect of time constant factors common to countries in specific world continents via regional fixed effects for Africa, America, Europe, Asia and Oceania (all regressions in Table 3).

We, moreover, control for the size of the public sector based on general government final consumption expenditure as percentage of GDP (regressions 6 to 9 in Table 3). This allows us to account

Table 3: Robustness check for the determinants of female participation rates estimation

Regressors:	OLS(5)	OLS(6)	OLS(7)	OLS(8)	OLS(9)
Maternity leave (λ)	0.563*	0.599*	0.819***	1.010***	1.037***
	(0.348)	(0.354)	(0.321)	(0.366)	(0.370)
Maternity leave squared(λ^2)	-0.011**	-0.011**	-0.015***	-0.017***	-0.016***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
$\lambda \times$ low GDP per capita (Dy)	-	-	-	-	0.008
	-	-	-	-	(0.207)
$\lambda^2 \times$ low GDP per capita (Dy)	-	-	-	-	-0.004
	-	-	-	-	(0.004)
Maternity leave threshold (weeks)	25.6	27.2	27.3	29.8	32.4
Abortion index (Λ)	1.766***	1.788***	1.439***	1.635***	1.648***
	(0.294)	(0.301)	(0.318)	(0.339)	(0.340)
Attitudes working women (Σ)	0.785***	0.816***	0.828***	0.677***	0.685***
	(0.078)	(0.082)	(0.084)	(0.100)	(0.100)
Maternity replacement rate (ρ)	0.022	0.025	0.040	0.037	0.036
	(0.029)	(0.031)	(0.031)	(0.032)	(0.032)
Infant mortality rate (m)	0.164***	0.145**	0.120**	0.119**	0.114**
	(0.054)	(0.059)	(0.057)	(0.058)	(0.058)
Urban population rate (u)	-0.201***	-0.199***	-0.143**	-0.117*	-0.108*
	(0.062)	(0.063)	(0.058)	(0.060)	(0.061)
Years of schooling male (h_m)	-1.675*	-1.918**	-1.733*	-2.242**	-2.331**
	(0.885)	(0.928)	(0.972)	(1.007)	(1.017)
Years of schooling female (h_f)	3.882***	3.971***	2.986***	3.328***	3.348***
	(0.872)	(0.916)	(0.961)	(0.949)	(0.950)
log GDP per capita ($log(y)$)	-54.409***	-56.625***	-55.886***	-52.621***	-51.523***
	(9.412)	(10.037)	(10.373)	(10.847)	(10.849)
log GDP per capita squared ($log(y)^2$)	2.722***	2.853***	2.818***	2.641***	2.561***
	(0.494)	(0.530)	(0.562)	(0.587)	(0.591)
Government consumption (% of GDP)	-	0.127	-0.146	-0.180	-0.200
	-	(0.157)	(0.163)	(0.171)	(0.174)
Time fixed effects	yes	yes	yes	yes	yes
Continental fixed effects	yes	yes	yes	yes	yes
Legal origins	no	no	yes	yes	yes
Religion	no	no	no	yes	yes
Number of observations	322	309	306	303	303
Adjusted R-squared	0.680	0.671	0.700	0.710	0.711

Note: *, **, *** measures statistical significance at the 10, 5 and 1 percent levels respectively. All regressions include a constant, fixed time trends and regional dummies for Asia, Africa, Europe and America and report robust errors. Final consumption of the general government appears in regressions (6)-(9). Regressions (7)-(9) include Legal origins, which are four dummy variables that identify the legal origin of the company law or commercial code of each country: (i) French commercial code; (ii) German commercial code; (iii) Scandinavian commercial code; (iv) socialist communist laws. Regression (9) also includes four dummies for religious affiliation (Protestants, Catholic, Eastern Orthodox, Muslim) as a percentage of population in 2000. Regression (9) includes the interaction between maternity leave and its square with a dummy variable Dy for countries with a GDP per capita lower than 15,500 constant PPS dollars.

somewhat for differences in the welfare states across countries and over time; differences that are likely to be associated with a range of model variables including female labor supply, maternity leave, fertility and social norms. In addition, (in regressions 7 to 9) we control for a country's legal tradition on the premise that this captures the preference for public intervention in the private sphere ranging from a relatively laissez-fair approach in common law systems (originating in the UK), greater intervention in civil law regimes (radiating from France, Germany and Scandinavia) and greatest intervention in countries with a socialist-communist legal heritage (La Porta et al., 2008). We also control for the percentage of the population belonging to the main monotheistic religions (Protestants, Catholics, Eastern Orthodox or Muslim) in the year 2000 (North et al. 2013). This allows us to consider the confounding effects of cultural norms beyond those that we focus on this article. As can be seen in regressions 7 and 8 of Table 3, the estimated effect of positive attitudes towards working women is reduced from 0.828 to 0.677. Although not shown, this is mostly due to the fact that we are now controlling for the percentage of the population that identify themselves as muslims. This religion seems to have a wider impact on female labor supply, above and beyond the social norms we focus on this article.

In addition, regression (9) allows for different slopes in maternity leave in low compared to high income countries by multiplying λ and λ^2 with a dummy variable Dy that takes the value of one when the country has a GDP per capita lower than the average value observed in the data (15,500 PPS dollars). Thus, the coefficients of $\lambda \times Dy$ and $\lambda^2 \times Dy$ measure the difference in the effect of maternity leave and its square between low and high income countries. These coefficients are not statistically different from zero, meaning that maternity leave duration has a similar effect in both types of countries.

Finally, figure 1 shows the predicted relationship between maternity leave and the female labor force participation rate obtained from the most saturated estimated equation (OLS(8) in Table 3). It clearly maps the inverted U-shape relationship between these two variables. The predicted female labor force participation rate first increases from 46.0% to a maximum of 60.3% when maternity leave goes from 1 week (the lowest observed value of maternity leave) to 30 weeks, and then decreases until it reaches a minimum of 45.7% when maternity leave is equal to 60 weeks (around the highest value of maternity leave in the data).

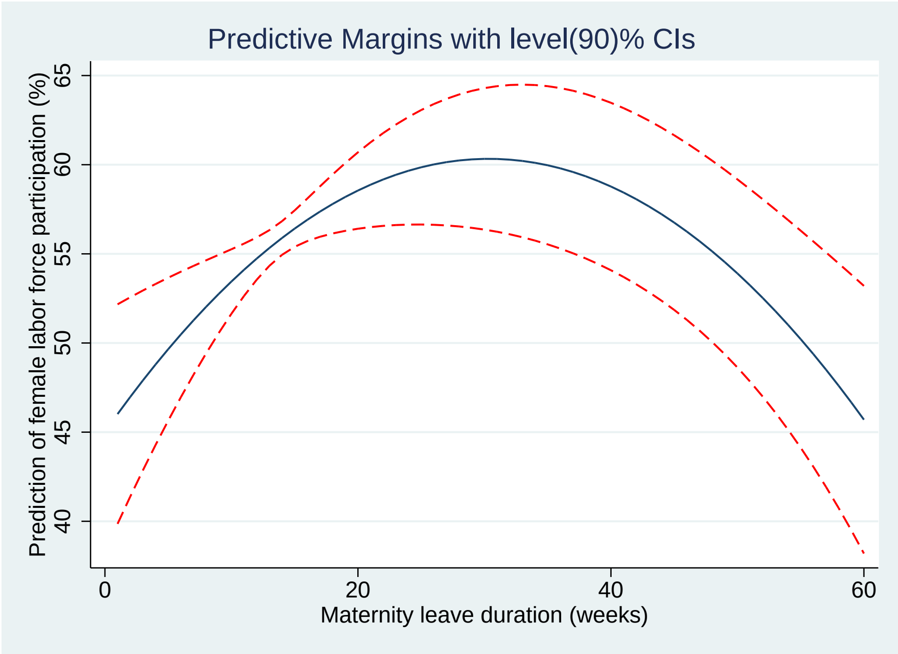

Figure 1: Predicted relationship between maternity leave and female labor force rate

Note: Predicted female labor force rate using regression OLS(8) in Table 3 with confidence intervals of 90%.

Before concluding, it is necessary to address the issue of reverse causality. As mentioned by Olivetti and Petrongolo (2017), changes in female labor supply may create political support for parental leave rights and lead to extended rights including additional weeks of maternity leave. Unaccounted for, reverse causality may potentially introduce upward bias in the estimated impact of maternity leave on female labor force participation (since stronger political support may reasonably be expected to increase leave duration). One way to address reverse causality would be by way of instrumental variables and 2SLS regressions. Unfortunately, valid instrumental variables are difficult to find, especially for the square of maternity leave duration. On the other hand, the impact of reverse causality is likely to be reduced by the inclusion of public sector size, legal origins and social norms in the regressions. Public sector size and legal origins help account for differences in the way political support for parental leave may lead to changes in public policies. Female labor market participation may also impact on maternity leave policies by changing social norms regarding female employment (Olivetti and Pentrongolo, 2017), thus highlighting the usefulness of including social norms in the regressions. Notwithstanding this discussion, the possibility persists for upward bias in the estimated effect of parental leave and therefore, the threshold maternity leave duration of 30 weeks reported should be taken as an upper bound.

7 Conclusion

It is generally acknowledged that the impact of maternity leave on labor market outcomes depends, in part, on the duration of maternity leave. Moreover, female labor market outcomes and fertility are clearly interconnected. We account both theoretically and empirically for the effects of paid and unpaid maternity leave duration on female labor force participation in the presence of fertility decisions.

From a theoretical perspective, we extend Bloom et al.'s (2009) unitary model by including maternity leave duration through two main channels: (i) the time cost of female market work and; (ii) women's earnings. The first channel increases female labor supply if the leave is fully paid. The wage penalty effect reduces labor supply because the absence from work, due to the birth of a child, does not only reduce a mother's earnings if the leave is not fully paid but may also have a negative effect on the marginal product of female labor. Due to the presence of on-leave related costs to the firm, we assume that the wage penalty associated with the productivity loss depends on the length of the leave awarded to mothers and that its effect is quadratic. Thus, our theoretical model considers non-monotonic effects of leave duration on labor supply including its indirect effect through fertility decisions. This is important because maternity leave duration induces higher fertility which has a negative effect on female labor supply. As a result, our model shows that the total effect of leave duration on female labor supply is ambiguous and it is thus entirely an empirical question.

We test the theoretical predictions of our model using an unbalanced panel covering the periods 1994, 2004 and 2011 and 159 countries. Our findings confirm the existence of an inverted U-shaped relationship between maternity leave duration and female labor force participation, and identify a maternity leave duration threshold of around 30 weeks above which female participation falls. According to the theoretical model, if one is on the left of this threshold, increasing maternity leave will increase female labor force participation because the positive effect due to the reduction of work-time cost of employed mothers strongly dominates the negative wage penalty effect. Alternatively, beyond this threshold, the latter effect will overwhelm the effect due to the reduction in the work-time cost. Previous studies have found a similar relationship but only in OECD economies. Thus, our study confirms the existence of an inverse U-shape effect of maternity leave duration on labor market outcomes for a much wider cross-section of countries. The 30 weeks threshold, however, should be taken as an upper bound due to potential presence of reverse causality from labor force participation to maternity leave entitlements.

We also study the effects of social norms on female labor force participation. Our estimates indicate that an increase of 10 percentage points (near to one standard deviation) in the prevalence of positive attitudes towards working women increases the female labor force participation by 6.8 percentage points. According to our theoretical model, the reduction in both the time cost of finding a job and the cost of reducing fertility may be behind this positive effect.

Finally, future research could explore the extent to which the non-linear effect of maternity leave duration may impact on other labor market outcomes, most notably, wages and gender gaps.

Compliance with Ethical Standards:

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Appendix

Following Bloom et al. (2009), in order to derive an equation to estimate, we linearize Eq. (7) around the point $(\bar{y}, \bar{h}_f, \bar{h}_m, \bar{\lambda}, \bar{\lambda}^2, \bar{\rho}, \bar{n}, \bar{m}, \bar{u}, \bar{\Sigma})$ and obtain:

$$
\begin{aligned}
l_f \approx\ & l_f(\bar{y}, \bar{h}_f, \bar{h}_m, \bar{\lambda}, \bar{\lambda}^2, \bar{\rho}, \bar{n}, \bar{m}, \bar{u}, \bar{\Sigma}) \\
&+ \left(-\frac{c_0 \alpha (1+\alpha)(1-\sigma)\bar{h}_f \left[1-\theta\bar{\lambda}^2\right]}{\left((1+\alpha)\left(1-\bar{\lambda}(1-\bar{\rho})\right)(1-\sigma)\bar{y}\bar{h}_f\left(1-\theta\bar{\lambda}^2\right)\right)^2} \right) * y \\
&+ \left(\frac{-\alpha}{(1+\alpha)\left(1-\bar{\lambda}(1-\bar{\rho})\right)\bar{h}_f\left(1-\theta\bar{\lambda}^2\right)} \right) * h_m \\
&+ \left(\frac{\alpha\left((1-\sigma)\bar{y}\bar{h}_m - c_0\right)}{(1+\alpha)\left(1-\bar{\lambda}(1-\bar{\rho})\right)(1-\sigma)\bar{y}\left(1-\theta\bar{\lambda}^2\right)\left(\bar{h}_f\right)^2} \right) * h_f \\
&+ \left(\frac{\lambda\alpha\left((1-\sigma)\bar{y}\bar{h}_m - c_0\right)}{(1+\alpha)(1-\sigma)\bar{y}\bar{h}_f\left[1-\theta\bar{\lambda}^2\right]\left(1-\bar{\lambda}(1-\bar{\rho})\right)^2} \right) * \rho \\
&+ \left(\frac{(1-b\bar{n}(1-\bar{m}))}{(1+\alpha)\left(1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma})\right)^2} - \frac{(1-\bar{\rho})\,\alpha\left((1-\sigma)y\bar{h}_m - c_0\right)}{(1+\alpha)(1-\sigma)\bar{y}\bar{h}_f\left(1-\bar{\lambda}(1-\bar{\rho})\right)^2\left(1-\theta\bar{\lambda}^2\right)^2} \right) * \lambda \\
&+ \left(-\frac{\theta\alpha\left((1-\sigma)y\bar{h}_m - c_0\right)}{(1+\alpha)\left(1-\bar{\lambda}(1-\bar{\rho})\right)(1-\sigma)\bar{y}\bar{h}_f\left(1-\theta\bar{\lambda}^2\right)^2} \right) * \lambda^2 \\
&+ \left(\frac{-b(1-\bar{m})}{(1+\alpha)\left[1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma})\right]} \right) * n + \left(\frac{b\bar{n}}{(1+\alpha)\left[1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma})\right]} \right) * m \\
&+ \left(\frac{-(1+\alpha)\,\psi'(\bar{\Sigma})}{\left((1+\alpha)(1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma}))\right)^2} \right) * \Sigma \\
&+ \left(-\frac{\phi(1-b\bar{n}(1-\bar{m}))}{(1+\alpha)\left[1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma})\right]^2} - \frac{\alpha(1+\alpha)\left(1-\bar{\lambda}(1-\bar{\rho})\right)c_0}{\left((1+\alpha)\left(1-\bar{\lambda}(1-\bar{\rho})\right)w_f(\bar{u})\right)^2}\frac{dw_f}{du} \right) * u \qquad \text{(A.1)}
\end{aligned}
$$

where $l_f(\bar{y}, \bar{h}_f, \bar{h}_m, \bar{\lambda}, \bar{\lambda}^2, \bar{n}, \bar{m}, \bar{u}, \sigma, \bar{\rho})$ collects all constant terms. Let β_x denote each term in brackets multiplying each variable $x = \{y, h_f, h_m, \lambda, \lambda^2, \rho, n, m, u, \Sigma\}$ in (A.1). Then, we can write the linearized equation (A.1) as (8). The predicted sign of each β can be readily seen in (A.1): $\beta_y < 0$, $\beta_{h_m} < 0$, β_{h_f} and β_ρ positive provided that $(1-\sigma)y\bar{h}_m - c_0 > 0$ (the substitution effect dominates, see footnote 7), $\beta_\lambda \lesseqgtr 0$, $\beta_{\lambda^2} < 0$, $\beta_n < 0$, $\beta_m > 0$, $\beta_\Sigma > 0$, $\beta_u < 0$.

We also linearize the first order condition determining optimal fertility given female labor supply

(9) around the point $(\bar{l}_f, \bar{\lambda}, \bar{m}, \bar{u}, \bar{\Sigma}, \bar{\Lambda})$ and obtain, collecting all constant terms in $n(\bar{l}_f, \bar{\lambda}, \bar{m}, \bar{u}, \bar{\Sigma}, \bar{\Lambda})$:

$$\begin{aligned} n \approx\; & n(\bar{l}_f, \bar{\lambda}, \bar{m}, \bar{u}, \bar{\Sigma}, \bar{\Lambda}) + \\ & \left(\frac{-(1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma}))}{b(1-\bar{m})} \right) * l_f + \left(\frac{\bar{l}_f}{b(1-\bar{m})} \right) * \lambda \\ & + \left(-\frac{\left[1-(1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma})\bar{l}_f\right]}{b\,(1-\bar{m})^2} + \frac{\eta\alpha}{\left(\eta(1-\bar{m})+k\left(\bar{\Sigma},\bar{\Lambda}\right)\right)^2} \right) * m \\ & + \left(\frac{-\phi\bar{l}_f}{b(1-\bar{m})} \right) * u + \left(\frac{-\psi'(\bar{\Sigma})\bar{l}_f}{b(1-\bar{m})} + \frac{\alpha k_\Sigma}{\left(\eta(1-\bar{m})+k\left(\bar{\Sigma},\bar{\Lambda}\right)\right)^2} \right) * \Sigma \\ & + \left(\frac{\alpha k_\Lambda}{\left(\eta(1-\bar{m})+k\left(\bar{\Sigma},\bar{\Lambda}\right)\right)^2} \right) * \Lambda \end{aligned} \tag{A.2}$$

that we can write as (10) letting γ_y denote each term in brackets multiplying each variable $y = \{l_f, \lambda, m, u, \Sigma, \Lambda\}$. The sign of each γ_y is clear from (A.2).

Substituting (10) into (8) and collecting terms we obtain

$$\begin{aligned} (1-\beta_n\gamma_{lf})\, l_f =\; & (\beta+\beta_n\gamma) + \beta_y y + \beta_{hm} h_m + \beta_{hf} h_f + \beta_\rho \rho + (\beta_\lambda+\beta_n\gamma_\lambda)\,\lambda + \beta_{\lambda^2}\lambda^2 \\ & + (\beta_m+\beta_n\gamma_m)\, m + (\beta_u+\beta_n\gamma_u)\, u + (\beta_\Sigma+\beta_n\gamma_\Sigma)\,\Sigma + \beta_n\gamma_\Lambda \end{aligned} \tag{A.3}$$

that we write as (11) letting μ_z denote each term in brackets multiplying variable $z = \{y, h_f, h_m, \lambda, \lambda^2, m, u, \Sigma, \Lambda\}$. From (A.3):

$$\begin{aligned} \mu_y &= \frac{\beta_y}{1-\beta_n\gamma_{lf}},\ \mu_{h_m} = \frac{\beta_{hm}}{1-\beta_n\gamma_{lf}},\ \mu_{h_f} = \frac{\beta_{hf}}{1-\beta_n\gamma_{lf}},\ \mu_\rho = \frac{\beta_\rho}{1-\beta_n\gamma_{lf}},\ \mu_\lambda = \frac{\beta_\lambda+\beta_n\gamma_\lambda}{1-\beta_n\gamma_{lf}}, \\ \mu_{\lambda^2} &= \frac{\beta_{\lambda^2}}{1-\beta_n\gamma_{lf}},\ \mu_m = \frac{\beta_m+\beta_n\gamma_m}{1-\beta_n\gamma_{lf}}, \mu_u = \frac{\beta_u+\beta_n\gamma_u}{1-\beta_n\gamma_{lf}}, \mu_\Sigma = \frac{\beta_\Sigma+\beta_n\gamma_\Sigma}{1-\beta_n\gamma_{lf}},\ \mu_\Lambda = \frac{\beta_n\gamma_\Lambda}{1-\beta_n\gamma_{lf}}. \end{aligned}$$

To determine the predicted signs of each μ_z, note first that the denominator is always positive:

$$1-\beta_n\gamma_{lf} = 1 - \frac{1}{(1+\alpha)} > 0$$

Therefore,

$$sign\ \mu_y = sign\ \beta_y < 0$$

and

$$sign\ \mu_{h_m} = sign\ \beta_{h_m} < 0.$$

The signs of μ_{h_f} and μ_ρ are, respectively equal to the signs of β_{h_f} and β_ρ, and positive in both cases if $(1-\sigma)y\bar{h}_m - c_0 > 0$. In turn,

$$sign\ \mu_\lambda = sign\ \ (\beta_\lambda+\beta_n\gamma_\lambda) \gtreqless 0,$$

$$sign\ \mu_{\lambda^2} = sign\ \beta_{\lambda^2} < 0,$$

and

$$sign\ \mu_m = sign\ \left(\beta_m + \beta_n \gamma_m\right) \gtrless 0.$$

Also,

$$sign\ \mu_u = sign\ \left(\beta_u + \beta_n \gamma_u\right) < 0$$

since, after some manipulation, and using $d = 1 - b\bar{n}(1-\bar{m}) - \bar{l}_f\left(1 - \bar{\lambda} + \phi\bar{u} + \psi(\Sigma)\right)$, and

$$\frac{dw_f}{du} = (1-\sigma)h_f\left(1-\theta\lambda^2\right)A'(u)(K/E)^\sigma > 0,$$

we can write,

$$\beta_u + \beta_n\gamma_u = -\frac{\phi d}{(1+\alpha)\left(1-\bar{\lambda}+\phi\bar{u}+\psi(\bar{\Sigma})\right)^2} - \frac{\alpha\,(1+\alpha)\left(1-\bar{\lambda}\,(1-\bar{\rho})\right)c_0}{\left((1+\alpha)\left(1-\bar{\lambda}\,(1-\bar{\rho})\right)w_f(\bar{u})\right)^2}\frac{dw_f}{du} < 0.$$

Finally,

$$sign\ \mu_\Sigma = sign\left(\beta_\Sigma + \beta_n\gamma_\Sigma\right) \lesseqgtr 0$$

and,

$$sign\ \mu_\Lambda = sign\left(\beta_n\gamma_\Lambda\right) > 0$$

since $\beta_n < 0$, and $\gamma_\Lambda < 0$.