An Accounting Exercise for the Shift in Life-Cycle Employment Profiles of Married Women Born Between 1940 and 1960

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Abstract

Life-cycle employment profiles of married women born between 1940 and 1960 shifted upwards and became flatter. We calibrate a dynamic life-cycle model of employment decisions of married women to assess the quantitative importance of three competing explanations of the change in employment profiles: the decrease and delay in fertility, the increase in relative wages of women to men, and the decline in child-care costs. We find that the decrease and delay in fertility and the decline in child-care cost affect employment very early in life, while increases in relative wages affect employment increasingly with age. Changes in relative wages,

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in particular returns to experience, account for the bulk (67 percent) of changes in life-cycle employment of married women.

1 Introduction

In the United States, as well as in many other developed countries, life-cycle employment profiles of married women born around mid-century changed in a noticeable way. Employment rates of women born in 1940 and earlier are low at *childbearing* ages (between age 20 to 35) and increase over the life-cycle. Changes in employment across cohorts are not uniform along the life-cycle, however. They are very pronounced at childbearing ages and more modest at later ages. As a result, life-cycle employment profiles of women born in 1960 not only shift upwards but also become much flatter.

In this paper, we build a dynamic life-cycle model of employment decisions of married women to assess the quantitative importance of three competing explanations of the change in life-cycle employment profiles: the decrease and delay in fertility, the increase in relative wages of women to men, and the decline in child-care costs. The incentives at work are not new. First, because child-rearing is intensive in women's time, employment at childbearing ages increases as fertility is reduced. Second, postponing fertility allows women to reach childbearing ages with a higher stock of accumulated work experience, thereby increasing their incentives to remain employed when having children. Finally, either an increase in women's wages relative to men or a decline in the cost of child-care makes working more attractive at childbearing ages, which feeds back on employment decisions later on in life because of experience accumulation.

After calibrating the model to the life-cycle facts characterizing the 1940 cohort, we show that the decrease and delay in fertility and the decline in child-care cost affect employment very early in life, while increases in relative wages affect employment increasingly with age. Assuming that the three forces account for 100 percent of the shift in life-cycle employment profiles, we find that changes in women's wages (in particular, returns to experience) account for 67 percent of the increase, versus 22 percent for cost of child-care, and 9 percent for fertility patterns (the residual term is equal to 2 percent). The effects of decrease and delay in fertility offset each other. Employment rates tend to increase following a decrease in fertility since reservation rates increase with the number of children. However, because of the presence of young children in the household, a delay in the timing of births tends to decrease fertility at later ages, since young children are more costly.

Our calibration procedure is new, as dynamic life-cycle models of employment decisions of married women are often estimated using maximum likelihood techniques (e.g., Eckstein and Wolpin 1989, Van der Klauuw 1996, or Francesconi 2002, to name only a few papers). Maximum likelihood is a more refined statistical procedure since it takes into account higher order moments, while we only match the average employment along the life-cycle.[1] Since large panel data sets are not available yet for the early cohorts we consider, we use a sequence of cross-sectional data from the Current Population Survey (CPS) from 1962 to 2004. Hence, we do not have all the information necessary to perform the maximum likelihood (i.e., conditional means and variances). We believe that calibrating the model is appropriate for the question at hand and find that it yields surprisingly good results. We obtain a very tight fit not only for the entire life-cycle employment profile of the 1940 cohort, but also for the employment by number of children at various ages. Moreover, we conduct sensitivity analysis to assess the robustness of our choice of parameter values.

The contribution of our accounting exercise is clear. Three influential papers have stressed the importance of changes in the pure gender wage gap (Jones, Manuelli, and McGrattan 2003), changes in returns to experience (Olivetti 2006), and changes in child-

[1] See Schoonbroodt (2002) and Eckstein and van der Berg (2005) for advantages and disadvantages of maximum likelihood versus moments estimation.

care costs relative to life-time earnings (Attanasio, Low, and Sanchez-Marcos 2004) to account for changes in women's labor supply either over time or across cohorts. Since our model nests these three potential explanations and adds another one (the decrease and delay in fertility), we can assess the quantitative importance of each of these forces separately. We find that they affect employment of women in distinct age groups differently and that changes in returns to experience have the largest impact on women's employment. Moreover, we show that a careful modeling of the distributions for number and timing of births is fruitful. First, it allows us to match the entire life-cycle employment of married women born in 1940. Second, once we control for changes in fertility patterns, exogenous changes in women's wages and cost of children that are needed to match changes in employment across cohorts are smaller in magnitude compared to the ones found in Jones, Manuelli, and McGrattan (2003) for the gender wage gap, Olivetti (2006) for returns to experience, and Attanasio, Low, and Sanchez-Marcos (2004) for decreases in the cost of child-care.

Numerous other explanations for the increase in employment of married women, either over time or across cohorts, have been proposed. These include falling prices of home appliances (Greenwood, Seshadri, and Yorukoglu 2005), changes in the perceived value of marriage (Caucutt, Guner, and Knowles 2002), the introduction of the pill (Goldin and Katz 2002), changes in social norms (Fernandes, Fogli, and Olivetti 2004), or gender-biased technological change favoring women (Galor and Weil 1996), to name only a few. These papers are certainly important. However, it is virtually impossible, let alone not desirable, to include all of the aforementioned forces into one single model. To perform our accounting exercise, we chose the ones which could be modeled without too much controversy and seemed the most likely to influence women's employment decisions at childbearing ages.

The paper is built as follows. In Section 2, we present evidence for the change in life-cycle patterns of employment and fertility for two cohorts of married women born in the United States in 1940 and 1960. In Section 3, we describe a dynamic life-cycle model of employment decisions of married women with experience accumulation. In Section 4, we explain our procedure for the calibration of the model. In Section 5, we perform the accounting exercise and, finally, we provide some concluding remarks in Section 6.

2 Data

We use data from the Current Population Survey (CPS) for the survey years 1964-2003 and from the decennial Census for the survey years 1970-2000 to describe the life-cycle patterns of employment and fertility for two cohorts of married women born in the United States in 1940 and 1960.[2]

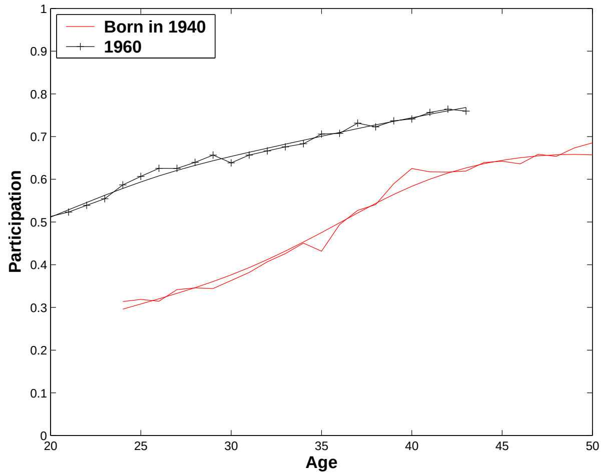
2.1 Employment

In Figure 1, we present the average employment by age for married women born in 1940 and 1960. We count as employed, any woman who was at work during the week preceding the interview or has a job but was not at work last week due to illness, vacations, etc. We pool data for women born within a three year interval (i.e., women born from 1939 to 1941 for the 1940 cohort and from 1959 to 1961 for the 1960 cohort) for the number of observations to be large enough at each age and we present both raw data as well as smoothed life-cycle employment profiles. Employment rates for women are low during childbearing ages (between age 20 to 35) and progressively increase over the life-cycle. Changes in employment rates across cohorts, however, differ in magnitudes along the life-cycle and are the largest at childbearing ages. Employment rates increased on average

[2]All raw data was downloaded from the Integrated Public Use Micro-data Series (IPUMS) available at http://www.ipums.org.

by 24 percentage points between age 20 and 35, compared to only 11 percentage points between age 36 and 50 (see Table 1). This fact is the focus of our analysis.[3]

Fig. 1: Life-Cycle Employment Profile of Married Women by Cohort



Tab. 1: Employment Rates of Married Women by Cohort and Age Group

	Age 20-35[a]	*Age 36-50*[b]	*Age 20-50*[c,d]
1940 Cohort	37	62	52
1960 Cohort	61	73	65
Change (in pct. points)	+24	+11	+13

[a]Age 24-35 for 1940 cohort. [b]Age 36-43 for 1960 cohort. [c]Age 24-50 for 1940 cohort. [d]Age 20-43 for 1960 cohort.

[3]In the Appendix, we show that increases in employment rates are the largest at childbearing ages throughout the education ladder. As a result, the increase in the fraction of women with a college degree can only account for a small fraction of the increase in women's employment across cohorts.

2.2 Fertility

We use Census data for the years between 1980 and 2000 to describe the distributions for the total number of children ever born and the age of mother at birth of first child of married women born in 1940 and 1960. We consider married women at age 40, assuming that fertility is close to completion at that age, and record the fraction with 0, 1,..., 4+ children, where 4+ denotes married women with at least 4 children. On average, women born in 1940 had 2.6 children by age 40, while those born in 1960 had 1.9 (see Table 2). Moreover, the decrease in the total number of children ever born mainly occurred from a redistribution of mass away from 3 and 4 children towards 0, 1, and 2 children (see Figure 2).

The age at birth of first child is not directly reported as part of the Census data. We use the age of the mother and the age of oldest child in the household to calculate a proxy for age of mother at birth of first child. For each number of children ever born, $f \in \{0, 1, 2, 3, 4+\}$, we record the fraction of women who have their first child at age $a \in \{20, 21, ..., 40\}$. On average, women born in 1940 had their first child at age 23, while those born in 1960 had their first child three and a half years later (see Table 2). This increase in the average can be decomposed into two components: first, the average age at birth of first child increased across cohorts for all levels of completed fertility (see Figure 3); second, women who have many children tend to have their first child early and the fraction of women with 0, 1, and 2 children increased (see Figures 2 and 3).[4]

[4]In the Appendix, we describe the total number of children ever born and age at birth of first child for women with different education. We find similar patterns, i.e. fertility levels declined and women have their first child later. However, changes in the total number of children ever born are the largest for High School graduates, while the delay in fertility is the largest for College graduates women.

Tab. 2: Fertility Levels and Timing of Births by Cohort - (Std. Dev.)

	Cohort 1940	Cohort 1960
Total Number of Children Ever Born :	2.6 (1.2)	1.9 (1.1)
Age of Mother at Birth of First Child:	23.2 (2.9)	26.7 (4.7)

Fig. 2: Completed Fertility by Cohort

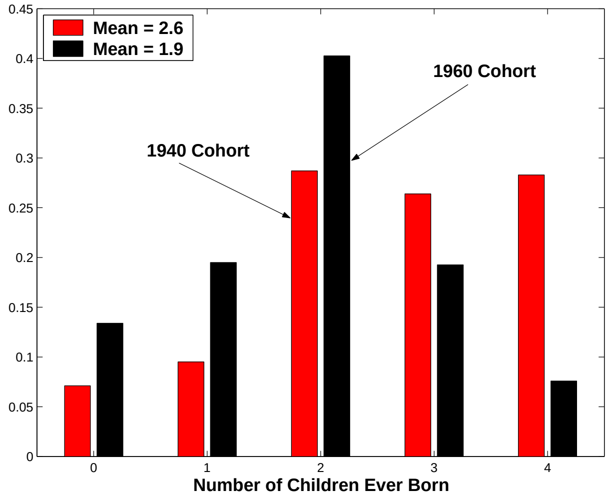
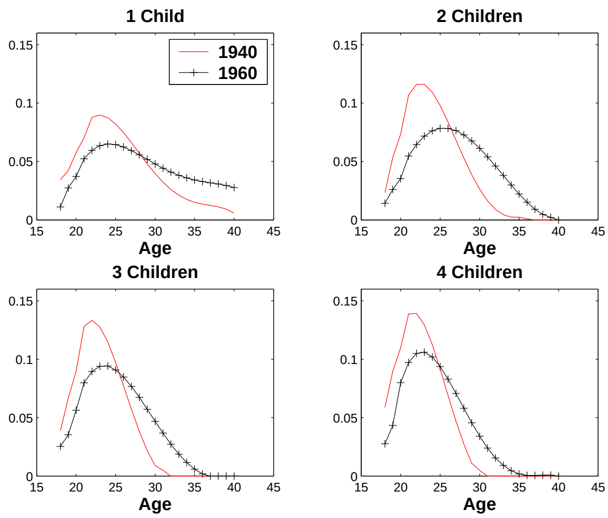


Fig. 3: Timing of Births by Completed Fertility and Cohort



2.3 Crossing Employment and Fertility

To understand how changes in the total number of children ever born and the age of mother at birth of first child affect employment rates along the life-cycle, we describe the employment decisions by number of children in the household at age 30 and 40 for our two cohorts (see Figures 4 and 5).

Focusing on the behavior of the 1940 cohort, it is clear that women's employment at age 30 is decreasing in the number of children in the household and that this effect is stronger for the first child. Note from Table 2 that the total number of children ever born decreased from 2.6 to 1.9 children per woman. Based on this fact alone, women's employment can increase across cohorts, due to a movement along a downward sloping curve. However, employment at age 30 also increased across cohorts for any given number of children. As childbirth is postponed, the fraction of women who used to have 2, 3, 4+ children at age 30 decreased and employment decreases with the number of children. Moreover, women born in 1960 are also more likely to have accumulated more work experience before childbearing, and therefore, are less likely to drop out of labor markets when having children. As a result, changes in the timing of births can account for the upward shift of the employment curve across cohorts. To assess the latter effect, a model of employment and experience accumulation is needed.

Finally, we present employment rates by total number of children ever born for married women at age 40 in Figure 5.[5] We find that, at least qualitatively, employment at age 40 also decreases with the number of children and that it increased across cohorts for any number of children. However, quantitatively, the impact of children on employment is not as strong as the one at age 30, as the differential in employment rates between women with no children and women with 4 children is much smaller than the same difference for

[5]We assume that women are no longer fertile after age 40 and present employment by number of children ever born rather than employment by number of children in the household.

Fig. 4: Employment of Married Women at Age 30 by Number of Children and Cohort

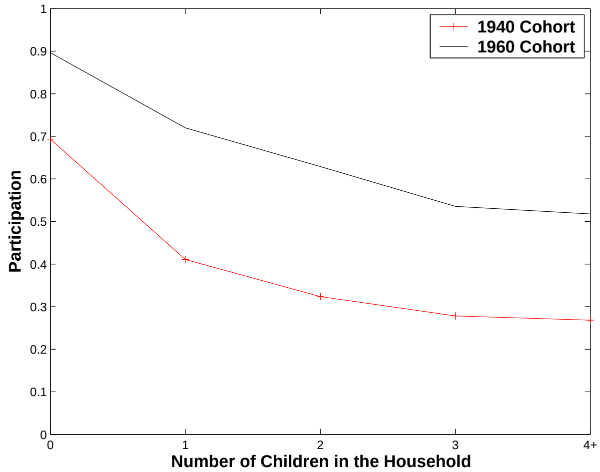
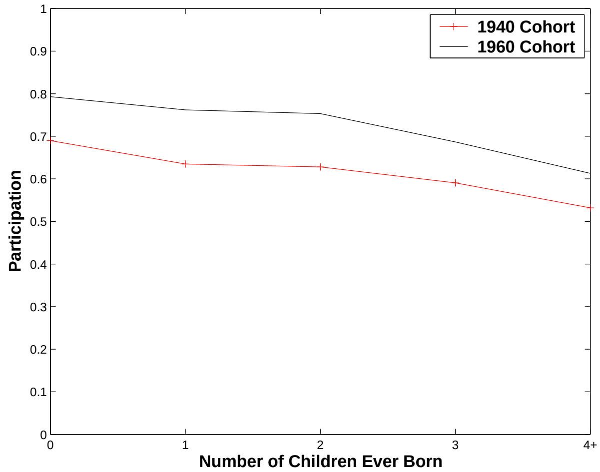


Fig. 5: Employment of Married Women at Age 40 by Number of Children and Cohort



women at age 30.

3 A Life-Cycle Model

In this section, we build the aforementioned economic mechanisms into a life-cycle model of employment decisions of married women with heterogenous agents and experience accumulation. Our model is close to Eckstein and Wolpin (1989).

3.1 Household's Maximization Problem

Demographics and Fertility: Men and women live with certainty for T periods and women are fertile for $T_f < T$ periods. Fertility is exogenous and women differ in the total number of children they have in a life-time, $f \in \{0, 1, ..., f_{max}\}$, and in the age at which they have their first child, a.[6] We fix the spacing of births to 2 years, so that the timing of all births is fully characterized by women's permanent type, (f, a): women of type $f \geq 1$ can have their first child by age $T_f - 2(f - 1)$ at the latest. Women know their type with certainty at the beginning of their life.

Preferences: Households derive utility from market consumption, c_t, and leisure time, l_t. We assume that the period-t utility, $U(c_t, l_t)$, is twice-continuously differentiable, increasing, and concave in both arguments, c_t and l_t.

Dynamic Optimization Problem: We model employment decisions of married women as a discrete choice, $e_t \in \{0, 1\}$.[7] At each age $t \in \{1, 2, ..., T\}$, women receive a wage offer, $w_t(h_t, \epsilon_t)$, which depends positively on work experience accumulated up to period t, h_t, and a contemporaneous productivity shock, ϵ_t. Women who accept the wage offer, i.e. $e_t = 1$, devote a fixed fraction of her time, $t_w \in (0, 1)$, to market activities and gain an additional year of work experience. The law of motion of work experience and women's wage offers are given by:

$$h_{t+1} = h_t + e_t \tag{1}$$

[6] Heckman and Walker (1990) find that the strongest effect of wages and costs of children operate through the time of the first birth.

[7] Since changes in women's labor supply across cohorts mainly occur at the extensive margin, this assumption is fine as a starting point. However, recent work by Erosa, Fuster, and Restuccia (2005) shows that, among working women, those who have children work fewer hours than the ones without children. Alternatively, Francesconi (2002) proposes a life-cycle model of women's labor supply and fertility where women can choose between working part-time or full-time. He finds that mothers prefer to interrupt their careers for a short time around childbirth rather than working on a part-time basis.

and

$$ln(w_t(h_t, \epsilon_t)) = \beta_0 + \beta_1 h_t + \beta_2 h_t^2 + \epsilon_t \tag{2}$$

where ϵ_t is normally distributed with mean 0 and standard deviation, σ_ϵ^2, and is i.i.d. over time.[8] We do not model joint participation decisions between husbands and wives. Men work with certainty in each period and their (deterministic) wage in period t is equal to w_{mt}.[9] Given the time discount factor, $\delta \in (0,1)$, women of type (f, a) choose employment, e_t, to maximize the expected discounted utility, $E_{t-1} \sum_{s=t}^{T} \delta^{s-t} U(c_s, l_s)$, subject to a sequence of budget and time constraints and the law of motion for work experience. In period t, the budget and time constraints are given by:

$$\begin{aligned} & c_t + g(f, a, e_t) \leq w_{mt} + w_t(h_t, \epsilon_t) e_t \\ & l_t + e_t t_w + t(f, a, e_t) = 1 \\ & e_t \in \{0, 1\} \end{aligned} \tag{3}$$

where the time-invariant functions, $g(\cdot,\cdot,\cdot)$ and $t(\cdot,\cdot,\cdot)$, denote the goods and time cost of children, respectively. Notice that we model the costs of children carefully, allowing them to depend on the age of children and women's participation choices. Following the work of Hotz and Miller (1988), we assume that both functions are increasing in the number of children and decreasing in age of children. On the other hand, goods costs increase

[8]The i.i.d assumption considerably reduces the dimension of the state space since we only need to keep track the current productivity shock as opposed to the entire history of shocks. In recent work, Meghir and Pistaferri (2004) and Guvenen (2005) reject the hypothesis that men's wage shocks are i.i.d over time and find strong empirical support for permanent and transitory wage shocks. However, since work experience is endogenous in our model, women's wages are serially correlated across periods even though productivity shocks are i.i.d.. Note that, if the woman works every period, work experience coincides with age and equation (2) boils down to a simple Mincer equation.

[9]Husband's wages are realized only after women's participation is made in Eckstein and Wolpin (1989) or Van der Klaauw (1996). Since they assume that utility is linear in consumption, women's participation decisions depend on husband's expected income.

with participation, while time costs decrease. This reflects the necessity of some sort of child-care when the woman works.

Our model abstracts from three important features. First, households cannot borrow or lend, implying that the only way to smooth consumption over the life-cycle is through women's labor supply.[10] Second, there is no depreciation in skills when women drop out of labor markets and only the stock of accumulated work experience, as opposed to the entire history of past employment decisions, matters to determine the average wage offers. Although these assumptions considerably reduce the dimension for the state space, Altug and Miller (1998) show that recent work experience is more valuable than distant one to determine women's wage offers. Finally, there are no permanent differences in women's market ability (fixed effects). Francesconi (2002) and Heckman and Walker (1990) find that high ability women are more likely to postpone fertility. Similarly, Van der Klaauw (1996) and Caucutt, Guner, and Knowles (2002) show that women with high market ability tend to postpone marriage (they wait for a suitable match), which, in turn, influences the age at which they have their first child and their employment decisions along the life-cycle. We briefly address this issue in Section 4.2.

3.2 Dynamic Program

We denote by $V_t(h, \epsilon; \theta)$ the maximum expected life-time utility discounted back to period t for women of type $\theta = (f, a)$, who are in state (h, ϵ). The household maximization problem can be formulated as a dynamic program, whose Bellman equation is given by:

$$V_t(h, \epsilon; \theta) = \max_{e_t \in (0,1)} \Big\{ U(c, l) + \delta E_t V_{t+1}\big(h', \epsilon'; \theta\big) \Big\} \tag{4}$$

[10] Attanasio, Low, and Sanchez-Marcos (2004) study a life-cycle model of women's employment with borrowing and savings. They show that the elasticity of women's employment increases once savings and borrowing are allowed.

subject to the law of motion (1), the earnings equation (2), and the budget and time constraints (3). Plugging the budget and time constraints into women's utility, we define the function, $W_t^{e_t}(h, \theta, \epsilon)$, as:

$$\begin{aligned} W_t^{e_t}(h, \theta, \epsilon) &= U(w_{mt} + w_t(h, \epsilon)e_t - g(f, a, e_t), 1 - e_t t_w - t(f, a, e_t)) \\ &+ \delta E_t V_{t+1}\big(h + e_t, \epsilon', \theta\big) \end{aligned} \tag{5}$$

Notice that W_t^0 is independent of ϵ_t, while W_t^1 is an increasing concave function of ϵ_t. As a result, there exists a reservation productivity shock, $\epsilon^*(h, \beta_i, \theta)$, such that women are indifferent between working and not-working, i.e. $W_t^0(h, \theta) = W_t^1(h, \theta, \epsilon_t^*)$, and women work if and only if $\epsilon_t \geq \epsilon_t^*(h, \beta_i, \theta)$.[11] In the Appendix, we derive the comparative statics of the productivity threshold. We show that, holding everything else the same, it decreases with work experience and the coefficients of Mincer wage equation, while it increases with the total number of children. As a result, life-cycle employment rates unambiguously increase following a left-shift in the distribution of total number of children ever born, or an increase in the coefficients of the Mincer wage equation, $(\beta_0, \beta_1, \beta_2)$. A shift in the distribution towards delay in fertility increases employment early on. However, there are two counterbalancing effects for later ages: (1) women born in 1960 are more likely to work since they have accumulated more work experience, (2) they are less likely to work since eventually they will have younger (i.e. more costly) children.

We solve the dynamic program using a standard backward induction procedure, assuming that the continuation value in period $T + 1$ is a function of work experience, $V_{T+1}(h)$. Given the expression for ϵ_t^*, the expected utility at time $t - 1$ is equal to:

$$E_{t-1}V_t(h, \theta) = \Phi\Big(\epsilon_t^*(h, \theta)\Big)W_t^0(h, \theta) + \int_{\epsilon_t^*(h,\theta)} W_t^1(h, \theta, \epsilon)\phi(\epsilon)d\epsilon \tag{6}$$

where ϕ and Φ denote the probability density function and the normal cumulative distribution for the productivity shocks. We use the functions ϵ_t and $E_t V_{t+1}$ to calculate the

[11]Note that, because of the i.i.d. assumption, the contemporaneous productivity shock enters the expression in (5) only once, through the woman's wage offer.

aggregate employment rates over the life-cycle in three steps. First, since women work when the productivity shock is higher than the reservation productivity, the average employment for women of type θ is equal to:

$$p_t(h,\theta) = 1 - \Phi(\epsilon^*(h,\theta)) \tag{7}$$

Second, we calculate the fraction of women, $\mu_t(h,\theta)$, of type θ who have accumulated h years of work experience at the beginning of period t. It is given by the following formula:[12]

$$\mu_{t+1}(h,\theta) = \mu_t(h,\theta)\big(1 - p_t(h,\theta)\big) + \mu_t(h-1,\theta)p_t(h-1,\theta) \tag{8}$$

with initial condition $\mu_1(0,\theta) = 1$ and $\mu_1(h,\theta) = 0$ for $h > 0$. Finally, the aggregate employment rate of married women in period t is equal to:

$$P_t = \sum_{(h,\theta)} \varphi(\theta)\mu_t(h,\theta)p_t(h,\theta) \tag{9}$$

where $\varphi(\theta)$ denotes the distribution over fertility types.

4 Calibration: 1940 Birth Cohort

In this section, we calibrate our model to the life-cycle facts characterizing the 1940 cohort.[13] We stress the importance of the distributions for the number and timing of births presented in the data section. Although dynamic discrete choice life-cycle models are usually estimated using maximum likelihood techniques (e.g., Eckstein and Wolpin 1989,

[12] The law of motion for μ is given by: $\mu_{t+1}(h,\theta) = \mu_t(h,\theta)\big(1 - p_t(h,\theta)\big)$ for women who have no prior work experience, i.e. $h = 0$. On the other hand, it is equal to $\mu_{t+1}(h,\theta) = \mu_t(h-1,\theta)p_t(h-1,\theta)$ for women who have worked in all periods, i.e. $h = t$.

[13] The calibration tool was introduced by Prescott (1986) and Kydland and Prescott (1982). It is now widely used in macroeconomics to assess the quantitative importance of dynamic general equilibrium model. Hansen and Heckman (1996) examine the empirical foundations of calibration.

Van der Klauuw 1996, or Francesconi 2002), the calibration yields surprisingly good results. We obtain a very tight fit not only for the entire life-cycle employment profile of the 1940 cohort, but also for the employment by number of children at various ages.

4.1 Parameter Values

1. *Demographics & Fertility*: The model period is one year. We consider women between age 20 to 60, i.e. $T = 41$. We assume that women are fertile between age 20 to 40, so $T_f = 21$. We set the maximum number of children, $f_{max} = 4$, so that women can have $f \in \{0, 1, 2, 3, 4\}$ children. We characterize the joint distribution $\varphi(\theta)$ in equation (9) using the distributions of number and timing of births for the 1940 cohort. Let $\varphi_f^{1940}(f)$ the marginal distribution of total number of children ever born as presented in Figure 2 of the data section and $\varphi_{a|f}^{1940}(a)$ the conditional distribution of the age of mother at birth of first child as presented in Figure 3. Then, the joint distribution in equation (9) is equal to: $\varphi(\theta) = \varphi_f^{1940}(f)\varphi_{a|f}^{1940}(a)$.

2. *Preferences*: Agents' utility is separable between consumption and leisure and is of the constant relative risk aversion form (CRRA). The period-t utility is given by:

$$U(c_t, l_t) = \frac{(c_t)^{1-\sigma_c} - 1}{1 - \sigma_c} + A\frac{(l_t)^{1-\sigma_l} - 1}{1 - \sigma_l} \tag{10}$$

for all values of σ_c and σ_l different from 1 and

$$U(c_t, l_t) = \ln(c_t) + A\ln(l_t) \tag{11}$$

when $\sigma_c = \sigma_l = 1$. A is a positive constant. Following Keane and Wolpin (2001) and Imai and Keane (2004), we set $\sigma_c = 0.52$, which implies a high value for the intertempotal elasticity of substitution in consumption (IESC) compared to previous studies.[14] They find that the introduction of *borrowing constraints* in life-cycle models significantly increase the value for IESC. We set $\sigma_l = 1$, following the

[14]With CRRA utility, the intertemporal elasticity of substitution in consumption (IESC) is equal to

indivisible labor supply model of Hansen (1985).[15] Sensitivity analysis shows that the model predictions crucially depend on the value of σ_c and σ_l.

3. *Costs of Children*: The goods and time cost of children functions, $g(f, a, e_s)$ and $t(f, a, e_s)$, are given by:

$$\frac{g(f,a,e_s)}{w_{mt}} = g_1 \mathrm{N}_s(f,a)^{\eta} + g_2 e_s \sum_{i=1}^{\mathrm{N}_s(f,a)} \rho^{s-a_i}$$
$$t(f,a,e_s) = (t_1 + t_2(1-e_s)) \sum_{i=1}^{\mathrm{N}_s(f,a)} \rho^{s-a_i}, \tag{12}$$
$$\text{with } (g_1, g_2, t_1, t_2, \rho, \eta) \in (0,1)^6$$

where $\mathrm{N}_s(f, a)$ denotes the number of costly children in the household at time s and $a_i = a + 2(i - 1)$ denotes the age of the i^{th} child. Notice that the goods cost of children is expressed as a fraction of husband's income and includes a base cost, g_1 and an additional cost, g_2, when women work. We interpret the latter as market child-care costs that arise when women work and have to find someone else to look after their child. We experiment on this parameter in relation to Attanasio, Low, and Sanchez-Marcos (2004).

Since $\eta < 1$, there are economies of scale in the goods cost of children. Similarly, the time cost of children includes a base cost, t_1 as well as an additional cost, t_2,

the inverse of the coefficient of risk aversion, σ_c (see Kimball 1990). Hubbard, Skinner, and Zeldes (1994) survey the literature on life-cycle consumption, savings, and wealth accumulation and conclude that a conventional value for σ_c is equal to -3, which implies a value for IESC of $-\frac{1}{3}$. They do not consider, however, imperfection in capital markets.

[15]A high IESL value is typically used in the real business literature in order to generate labor supply volatility close to that of the US data (e.g., Kydland 1995). On the other hand, estimates from micro panel data suggest that the intertemporal elasticity of labor supply of men in their prime-age is close to 0 (see Altonji 1986 or MaCurdy 1981). Using lotteries, Hansen (1985) shows that the indivisible labor supply model generates a large inter-temporal elasticity of labor supply at the aggregate level despite the fact that hours worked conditional on being employed are constant.

when women do not work. Following Hotz and Miller (1988), we assume that the time costs of children decreases at rate, $\rho < 1$, when children grow. Finally, we assume that children are costly until age 13.

We use evidence and estimates from the micro-econometrics literature to calibrate the parameters for the costs of children: $(g_1, g_2, t_1, t_2, \rho, \eta)$. Our main reference is Hotz and Miller (1988) who use a structural life-cycle model to estimate the time and goods of children. First, they find that the time cost of children decreases at rate 0.89 with age of children. Accordingly, we fix $\rho = 0.89$. Second, we set $g_1 = 0.09$ and $g_2 = 0.07$. This is in the upper range of Hotz and Miller estimates, who find that the goods cost per child per week ranges from 11 to 17 percent of husband's income.[16] Third, we fix $t_1 = 0.10$ and $t_2 = 0.06$, which compares well to their estimates. They find that the time cost of a newborn is about 13 percent of a woman's time after sleeping and eating hours have been subtracted.[17] Finally, Lazear and Michael (1980) find large economies of scale, while Espenshade (1984) find that they are of the order of five percent for an additional child. We take an intermediate stand and fix $\eta = 0.92$.

4. *Discount factor*: We set $\delta = 0.96$ to match an annual interest rate of roughly 4%.

5. *Male Wages*: We calculate the average weekly wage by age for married men born in 1940.[18] Assuming that men participate in labor markets in all period with

[16]Note that it is very common to find wide ranges of goods cost estimates in the literature. See also Bernal (2004) who finds a comparable wide range for child-care expenditures.

[17]Hill and Stafford (1980) analyzing time use data in 1976 find that women spend 550 minutes per child per week in child-care if they have one preschooler and 440 minutes per child per week if they have two (p.237). This corresponds to about 10 percent of a woman's total time after sleeping and eating hours have been subtracted. However, housework time can to some extent be viewed as time spent where watching children is possible at the same time.

[18]The Current Population Survey (CPS) provides individual data on total labor income earned in the

probability one, we fit the average observed wage of men over the life-cycle using a polynomial equation of degree 4:

$$ln(w_{m,age}) = \beta_{0m} + \beta_{1m}age + \beta_{2m}age^2 + \beta_{3m}age^3 + \beta_{4m}age^4 \tag{13}$$

We find the following parameters values: $\beta_{0m} = 5.7083$, $\beta_{1m} = 0.0805$, $\beta_{2m} = -0.0042$, $\beta_{3m} = 0.0001$, $\beta_{4m} = -9.4218e^{-7}$.

6. *Workweek length*: From time-use data (see Juster and Stafford 1991), people use on average 8 hours a day for sleeping and 2 for eating which leaves 98 hours per week to devote to work, leisure,.... From CPS data, the average workweek length for married women (conditional on being employed) is 35 hours a week. Therefore, $t_w = 35/98 = 0.36$ (see Greenwood, Seshadri, and Yorukoglu 2005).

7. *Women's Wages, Terminal Condition, and Marginal Utility of Leisure*: We assume that the continuation value function in period $T+1$ depends on work experience and is of the following form: $V_{T+1}(h) = a_1 h^{a_2}$ with $a_1 > 0$ and $a_2 > 0$.

   For women's wages, we first use Guvenen (2005)'s estimates for the variance of the productivity shocks and fix $\sigma_\epsilon^2 = 0.061$. Second, due to non-random selection of married women into the labor market, the wage coefficients of the Mincer equation, $(\beta_0, \beta_1, \beta_2)$, are potentially biased.

   To address this problem, we choose women's wage coefficients, marginal utility of leisure, and parameters for the continuation value, i.e. $\psi = \{\beta_0, \beta_1, \beta_2, A, a_1, a_2\}$, to minimize the squared deviation between the life-cycle employment rates from the model, $\{P_t(\psi; \xi)\}_{t=24}^{50}$, and their data counterpart for the 1940 cohort, $\{P_t^{d,1940}\}_{t=24}^{50}$:

$$\mathrm{Q}^c(\psi; \xi^c) = \sum_t \Phi_{t,t}^{-1} (P_t(\psi; \xi^c) - P_t^{d,1940})^2 \tag{14}$$

previous calendar year as well as weeks worked last year. Weekly wages are then total labor income divided by weeks worked.

where the elements of the weighting matrix, Φ^{-1}, are equal to the variance of participation rates over the life-cycle on the diagonal and zero otherwise. The vector of calibrated parameters, ξ^c, is equal to: $\xi^c = \{\{\varphi_f^{1940}\}, \{\varphi_{a|f}^{1940}\}, \sigma_c, \sigma_l, g_1, g_2, t_1, t_2, \rho, \eta, \delta, \{\beta_i^m\}, t_w, \sigma_\epsilon\}$.[19]

Notice that the system in equation (14) is over-identified since we have 27 moments to determine 6 parameters. As a result, we cannot match all the moments perfectly. However, the fit between moments and data is good as the minimum distance for the quadratic form is equal to 0.01, i.e. $\mathrm{Q}^c(\psi^c;\xi^c) = 0.01$ (see Table 3). We find that $\beta_0 = 5.3117$, $\beta_1 = 0.0105$, and $\beta_2 = -2.04e^{-4}$. Previous studies also find the sign of β_1 and β_2 to be positive and negative. However, our estimates are smaller than estimates from traditional Mincer regressions.[20] Finally, $A = 21.65$, $a_1 = 1.19$ and $a_2 = 0.44$.

Tab. 3: Calibrated Wage Parameters

β_0	β_1	β_2	$\mathrm{Q}^c(\psi^c(\cdot);\xi^c(\cdot))$
5.3117	0.0105	-2.04e^{-4}	0.01

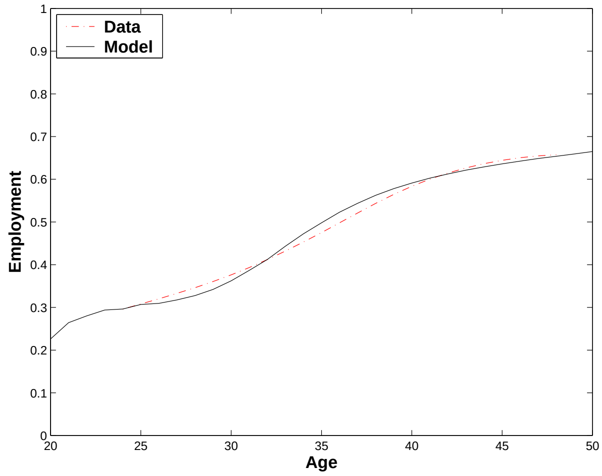
[19]We use the *downhill simplex* method to solve for the optimal vector, $\psi^c = \underset{\psi \in \Psi}{\arg\min} \mathrm{Q}^c(\psi;\xi^c)$, which requires only function evaluations, not derivatives, and is efficient when the size of the simplex is small (see Nelder and Mead 1965).

[20]This result is consistent with the findings of Eckstein and Wolpin (1989), who show that simple wage regressions on female wages yield biased estimates because of non-random selection in labor markets and experience accumulation. They find that, when using a structural model of women's employment decisions, the coefficient on experience and experience squared in the Mincer equation, β_1 and β_2, are equal to 0.0241 and $-2.4e^{-4}$, respectively, compared to 0.037 and $-5e^{-4}$ in simple wage regressions.

4.2 Cohort 1940: Model versus Data

In this section, we compare the model predictions versus data for calibrated moments as well as non-fitted moments. The calibrated life-cycle employment profile is quite close to the data (see Figure 6).

Fig. 6: Calibrated Life-Cycle Employment of Married Women - 1940 Cohort



We also explore other predictions of the model for moments that we did not calibrate directly. First, the model slightly over-predicts employment by number of children at age 30, while the fit is almost perfect at age 40 (see Figures 7 and 8, respectively).

Fig. 7: Employment at Age 30 by Total Number of Children Ever Born - 1940 Cohort

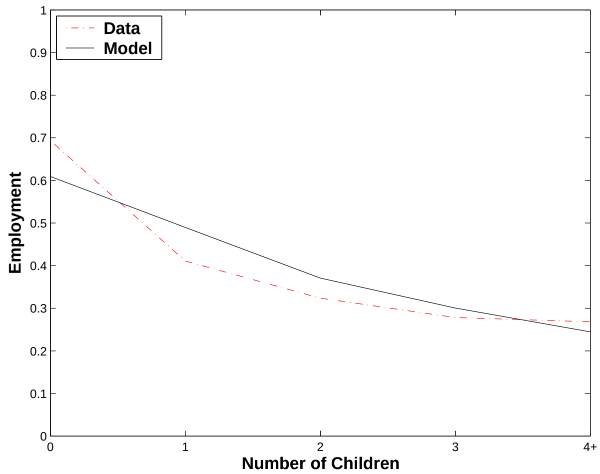
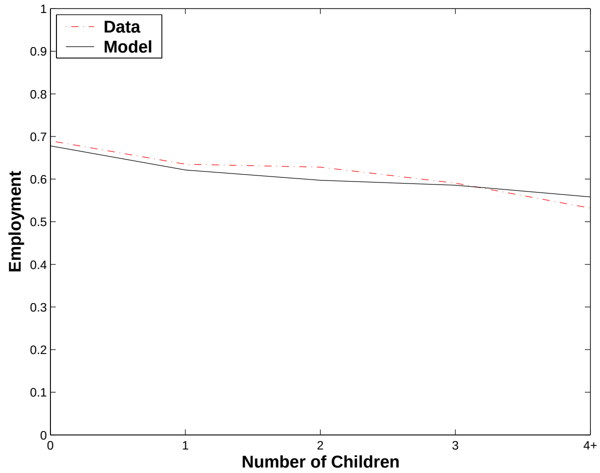


Fig. 8: Employment at Age 40 by Total Number of Children Ever Born - 1940 Cohort



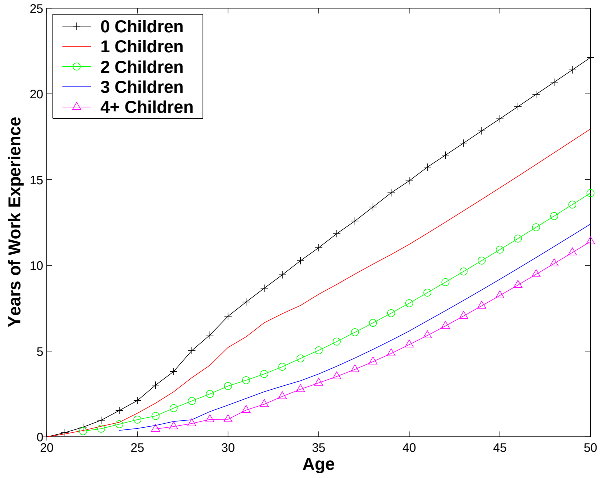
Second, since employment rates decrease with the number of children, women with fewer children tend to accumulate a greater number of years of work experience (see Figure 9). At age 20, women start with no work experience. By age 50, the experience gap between women who have no children and those who have 4+ children is greater than 11 years of work experience. All of the above findings suggest that shifts in the distribution of completed fertility (total number of children ever born in a life-time) as shown in Figure 2 in the data section potentially account for a large part of the increase in participation across cohorts. We quantify this statement in the next section.

We next address the model's predictions for the average observed wage over the life-cycle (see Figure 10) and the average observed wage by total number of children ever born at age 40 (see Figure 11).[21] Although we match the average wage over the life-cycle, the model overstates wages at early ages and fails to capture the increase in wages at later ages. Qualitatively, wages at age 40 decrease with the number of children. Quantitatively, however, children have a much smaller impact on wages than in the data.

Recall that in the present model, the reasons for wages to be decreasing in the number of children are (1) due to the relative goods and time costs of children, women with more

[21]We normalize wages by number of children by the wage of women with 0 children.

Fig. 9: Life-Cycle Years of Work Experience by Total Number of Children Ever Born - 1940 Cohort



children are less likely to work and hence accumulate less work experience, (2) women who delay childbirth are more likely to have accumulated more work experience before childbearing, and therefore, receive higher wage offers. This formulation misses out on an important dimension, namely that higher ability women (e.g., college educated women) tend to have children later and to have fewer. One way to account for this fact is to introduce fixed effects as an additional source of heterogeneity (i.e., market ability, β_0^i) and to allow for market ability to be positively correlated with age at birth of first child (which itself is negatively correlated with number of children). In such a model we find that the higher the correlation between market ability and age at birth of first child, the faster average wages fall as the number of children increases (addressing the the problem in Figure 11). The drawback of introducing fixed effects, however, is that changes in the distribution of ability types (i.e., wages levels) also change fertility related distributions, and vice versa. Hence counterfactual experiments such as those we perform in Section 5 are hard to interpret and call for arbitrary adjustments, unless independence is assumed. But under independence the aforementioned additional effect disappears. We therefore chose to use only one ability type, implicitly assuming independence.[22]

Fig. 10: Life-Cycle Weekly Wages of Married Women - 1940 Cohort

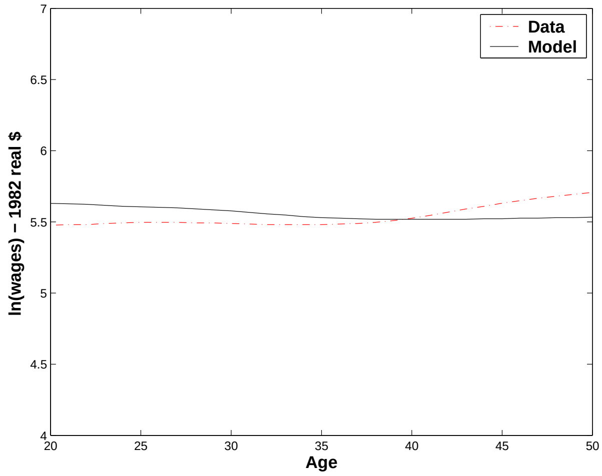
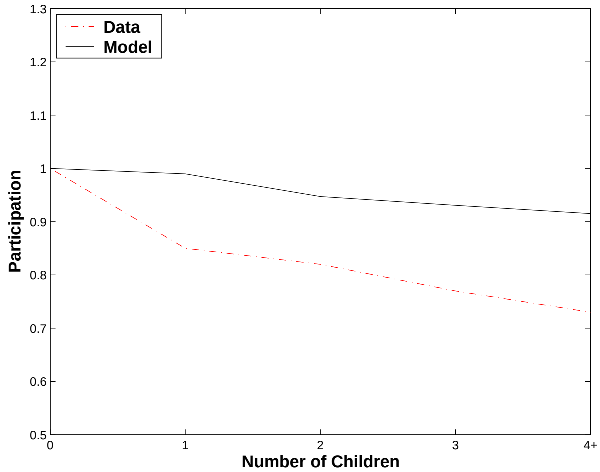


Fig. 11: Wages at Age 40 by Total Number of Children Ever Born - 1940 Cohort



[22]Introducing fixed effects becomes useful in a model with endogenous fertility, where high ability women *choose* to have fewer children and to have them later. As a result a change in wages (or the distribution of abilities) will affect both, fertility choices and participation decisions and no arbitrary adjustments are needed. This is the object of work in progress by the authors.

4.3 Sensitivity Analysis

In this section, we perform sensitivity analysis to assess the robustness of our calibrated parameters. We analyze the impact of a 10-percent change in our calibrated parameters on the goodness of fit for life-cycle participation and participation by number of children at age 30 and 40. Changing only 1 parameter at a time, we choose women's wages coefficients, the marginal utility of leisure, and parameters for the continuation value to minimize the squared deviation between the life-cycle employment rates from the model and their data counterpart. For example, to assess the impact of a 10-percent increase in the base goods cost, g_1, we set the vector, $\psi = \{\beta_0, \beta_1, \beta_2, A, a_1, a_2\}$, to minimize the following quadratic form:

$$\mathrm{Q}^s(\psi; \xi^s(g_1)) = \sum_t \Phi_{t,t}^{-1}(P_t(\psi; \xi^s(g_1)) - P_t^{d,1940})^2 \tag{15}$$

where $\xi^s(g_1) = \{\{\varphi_f^{1940}\}, \{\varphi_{a|f}^{1940}\}, \sigma_c, \sigma_l, 1.1g_1, g_2, t_1, t_2, \rho, \eta, \delta, \{\beta_i^m\}, t_w, \sigma_\epsilon\}$. Let $\psi^s(g_1) = \arg\min_{\psi \in \Psi} \mathrm{Q}^s(\psi; \xi^s(g_1))$, the solution to the above system. We calculate the percentage change (elasticity) in the coefficients of women's wages and marginal utility of leisure as follows: $\lambda_{\psi_i, g_1} = \frac{\psi_i^s(g_1) - \psi_i^c}{\psi_i^c} / \frac{\Delta g_1}{g_1}$. We assess the impact of a 10-percent change in parameters for preferences, cost of children, and the standard deviation of the productivity shock and present our results in Table 4.

Since the marginal utility of consumption for women with children is very high, the relationship between employment and number of children flattens out following an increase in the base goods cost of children, g_1 (relative to g_2).[23] The decrease in employment rates of women with 0 and 1 child is achieved through an increase in the marginal utility of leisure, A, and a decrease in returns to work experience, (β_1, β_2). On the other hand, when the base time cost t_1 is high relative to t_2, the marginal utility of leisure for women with children is very high. As a result, the relationship between employment and number

[23] For very high values of g_1 relative to g_2, the marginal utility of consumption for women with children is so high that employment rates increases with the number of children, which is counterfactual.

of children becomes steeper. Finally, the goodness of fit for life-cycle employment rates is very sensitive to values of ρ, which provides strong support for the estimate of Hotz and Miller (1988).

The wage-experience profile and the marginal utility of leisure are also very sensitive to the coefficient of risk aversion, σ_c and the intertemporal elasticity in labor supply, σ_l, which confirms the low values for σ_c in life-cycle models with borrowing constraints (see Keane and Wolpin 2001). Note that since the marginal utility of leisure decrease when σ_l increases, employment by number of children flattens out. Finally, changes in the standard deviation of the productivity shocks have mild effects on the goodness of fit for life-cycle participation and participation by number of children at age 30 and 40.

Tab. 4: Impact of 10% Increase in Calibrated Parameters

	β_0	β_1	β_2	A	$Q^s(\psi^s(\cdot);\xi^s(\cdot))$
g_1	-0.0011	-0.0538	-0.1353	0.0052	0.05
t_1	0.0006	-0.1811	-0.1142	-0.0019	0.03
ρ	0.0008	-0.1646	-0.3764	-0.0080	0.63
σ_c	0.2979	-1.0467	-0.2870	-0.1289	0.22
σ_l	-0.0942	-3.4774	3.2853	-0.6017	0.10
σ_ϵ	-0.0013	-0.2331	0.18568	0.0045	0.18

5 Experiments: 1960 Cohort

Three influential papers have stressed the importance of changes in the pure gender wage gap (Jones, Manuelli, and McGrattan 2003), changes in returns to experience (Olivetti 2006), and changes in child-care costs relative to life-time earnings (Attanasio, Low, and Sanchez-Marcos 2004) to account for changes in women's labor supply either over time or across cohorts. In this section, we assess the quantitative importance of these 3 forces as follows.

Taking changes in fertility patterns into account, we use our model to quantify what changes in women's wages and child-care cost are needed to match the life-cycle participation choices of women born in 1960. Using distributions for number and timing of births of the 1960 cohort, we choose the pure gender wage gap, β_0 (relative to β_{0m}), the returns to experience, β_1, β_2, and the cost of child-care, g_2, to match the life-cycle employment rates of the 1960 cohort, holding all other calibrated parameters constant. We set the vector, $\psi = \{\beta_0, \beta_1, \beta_2, g_2\}$, to minimize the following quadratic form:

$$\mathrm{Q}^e(\psi; \xi^e) = \sum_t \Phi_{t,t}^{-1} (P_t(\psi; \xi^e) - P_t^{d,1960})^2 \tag{16}$$

where $\xi^e = \{\{\varphi_f^{1960}\}, \{\varphi_{a|f}^{1960}\}, \sigma_c, \sigma_l, A, g_1, t_1, t_2, \rho, \eta, \delta, \{\beta_i^m\}, t_w, \sigma_\epsilon, a_1, a_2\}$.

Let $\psi^e = \arg\min_{\psi \in \Psi} \mathrm{Q}^e(\psi; \xi^e)$, the solution to the above system. This exercise allows us to answer questions such as: taking into account changes in fertility patterns, by how much do the coefficients of women's Mincer wage equation and child-care cost need to change to explain the observed patterns in women's employment?

Our model has the same qualitative predictions as in Jones, Manuelli, and McGrattan (2003), Olivetti (2006), or Attanasio, Low, and Sanchez-Marcos (2004). The pure gender wage gap and the child-care cost decrease, while returns to experience increase across cohorts in order to match changes in women's employment across cohorts (see Table 5). Quantitatively, however, we find that these changes are smaller than previously reported in these papers. The pure gender wage gap decreases, i.e. β_0 increases by less than 1 percent, returns to experience increase as the coefficient on experience, β_1, and experience squared, β_2, increase by 43 percent and 21 percent, respectively, and finally the price of child-care, g_2, decreases by 18 percent.

The fit for life-cycle employment profile of the 1960 cohort following changes in wage and fertility is good (see Figure 12). Since the increase in participation rates occurs mainly for women with 1 and 2 children, the experience gap between women with 0 and 1 child is less than 1.5 year at age 43 compared to more than 5 years for women born in 1940 (see Figure 15). Finally, employment at age 30 increases the most for women with 1 and 2 children and we overshoot for participation of women with 0 and 1 child at age 40 (see Figures 13 and 14).

Tab. 5: Wages & Child-Care Costs Needed to Match Life-Cycle Employment of 1960 Cohort

β_0	β_1	β_2	g_2	$Q^e(\psi^e(\cdot); \xi^e(\cdot))$
5.3217	0.0149	-1.61e^{-4}	0.057	0.06

Fig. 12: Employment over the Life-Cycle - 1960 Cohort

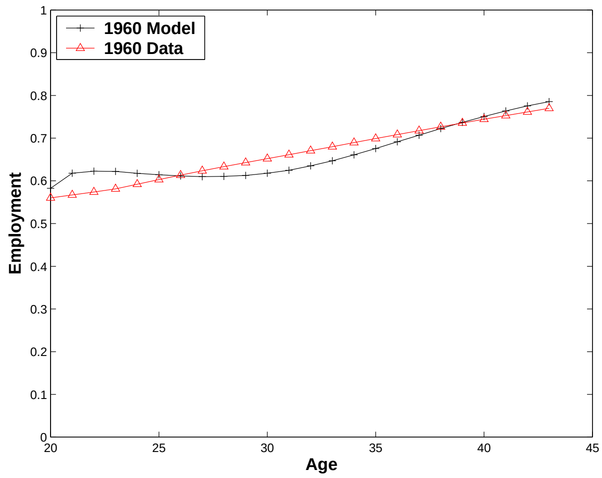


Fig. 13: Employment by Total Number of Children Ever Born at Age 30 - 1960 Cohort

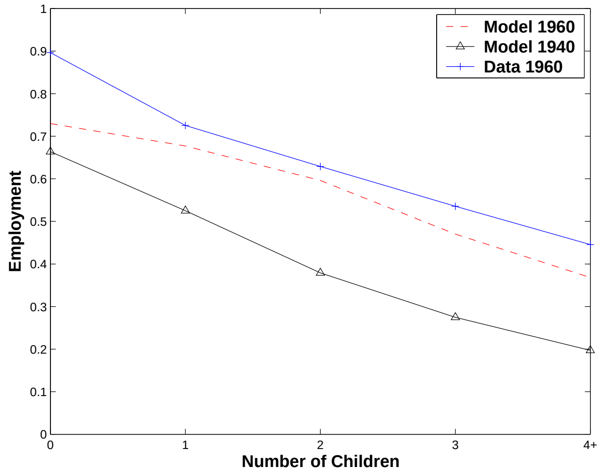


Fig. 14: Employment by Total Number of Children Ever Born at Age 40 - 1960 Cohort

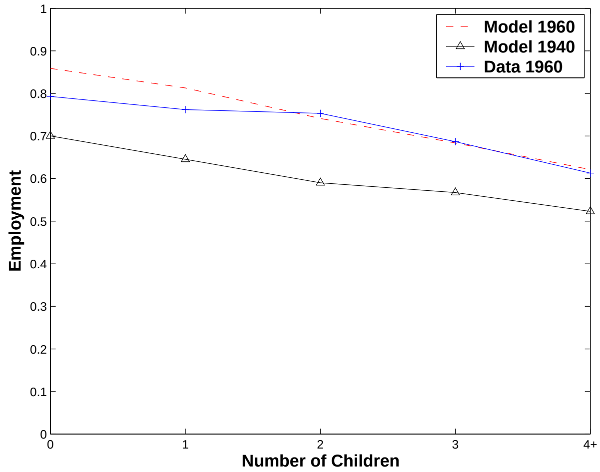
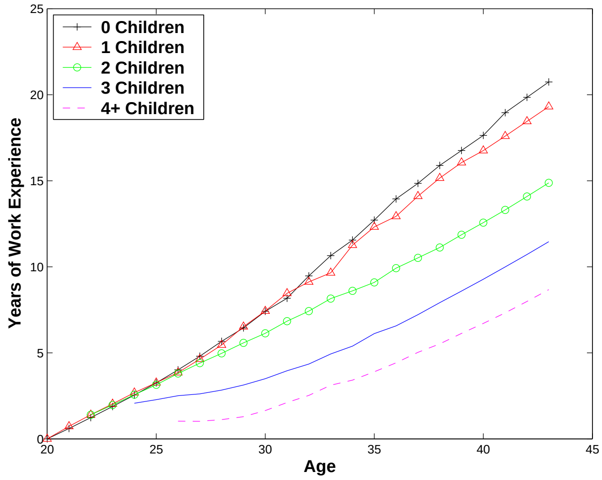


Fig. 15: Number of Years of Work Experience over the Life-Cycle by Total Number of Children Ever Born - 1960 Cohort



Since we implicitly assumed that changes in fertility patterns, gender wage differentials, and cost of child-care account for 100-percent of changes in women's employment across cohorts, we perform a decomposition exercise to assess their relative (quantitative) importance. We write:

$$1 = \Delta(\text{Fertility}) + \Delta(\text{Wages}) + \Delta(\text{Child-Care}) + R \tag{17}$$

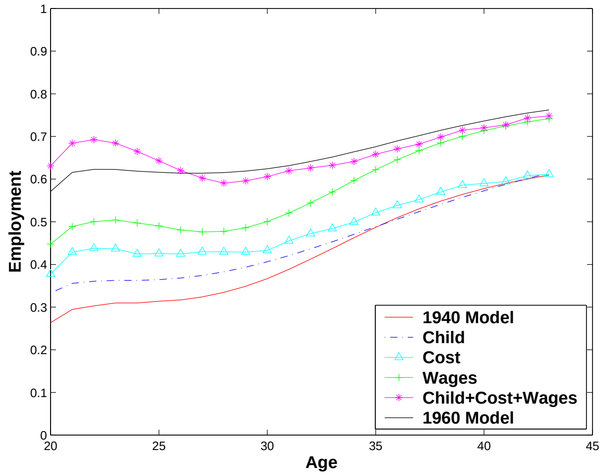
where changes in fertility include changes in number and timing of births, changes in wages include changes in the pure gender wage gap and returns to experience, and R is a residual term to account for potential interaction between all the variables. We present the results of the decomposition exercise for various age groups in Figure 16 and Table 6.

In order of importance, we find that changes in women's wages account for 67 percent of changes in women's employment across cohorts, followed by changes in child-care cost (22 percent), fertility (9 percent), and the residual term (2 percent). Overall, our results are in line with Olivetti (2006) and Caucutt, Guner, and Knowles (2002): changes in returns to experience have a large impact on changes in women's employment, while changes in the pure gender wage gap are more modest. Moreover, changes in fertility patterns have the largest impact for the age group 20-35 and changes in the number and timing of births offset one another. Finally, changes in fertility patterns also affect changes in women's life-cycle employment in an indirect way through their interaction with changes in women's wages and cost of child-care.

6 Concluding Remarks

We have presented data on employment and fertility for cohorts of married women born between 1940 and 1960. Using a life-cycle model of married female employment with experience accumulation, our analysis shows the following. First, fertility patterns (total number of children ever born and age of mother at birth of first child) are crucial de-

Fig. 16: Decomposing the effects of fertility, wages, and child-care costs



Tab. 6: Decomposing the effects of fertility, wages, and child-care costs (in percent)

	Age 20-35	*Age 36-43*	*Age 20-43*
Fertility:	15%	-6%	9%
Number of Children, f	20	12	17
Timing, a	-4	-18	-9
Wages:	57%	83%	68%
Pure Wage Gap, β_0	7	5	6
Returns to Experience, β_1	50	78	62
Child-Care Cost, g_2:	26%	14%	21%
Residual:	-8%	9%	2%
Fertility + Wages + Child-Care:	100%	100%	100%

terminants of the life-cycle employment profile for married women. Second, changes in gender wage differentials and costs of children needed to account for changes in women's life-cycle employment are smaller than previously found in the literature once we take into account changes in fertility patterns. Third, changes in women's wages (in particular, returns to experience) have the largest impact on women's employment decisions.

One open question is: What caused the decrease and delay in fertility? Our current work in progress is to endogenize fertility and timing of births decisions to ask whether changes in wages can also account for the delay in fertility. Preliminary results show that the delay, though positive, is largely left unaccounted for. Something else seems to have caused the delay in fertility. Several related questions come to mind: How is fertility related to the marriage decision? Was the change in fertility decisions simply due to cultural changes and changes in social norms? In what sense and can we attempt to measure these changes? The answers to these questions are closely intermingled and hard to disentangle from the question about why women tend to have children later and take care of them differently than they used to. They are however crucial to set up a useful model of fertility choices in terms of number and timing of births as well as child-care arrangements.

The other open question is why wage levels and returns to experience changed more for women than they did for men. Besides straight out discrimination, many other more readily quantifiable hypotheses can be considered. Changing occupational opportunities due to the rise in the service sector, changing educational investments pertaining more to women than men because of initial conditions are only a few avenues to be explored further.

7 Appendix

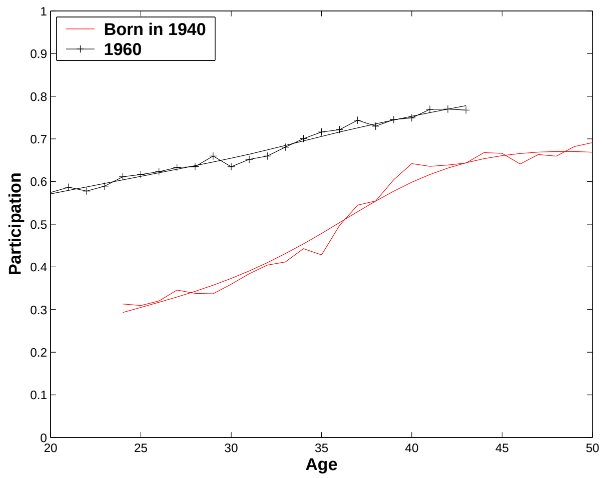
7.1 Life-cycle Patterns by Education

We present data for life-cycle employment profiles and fertility by education.

7.1.1 Employment

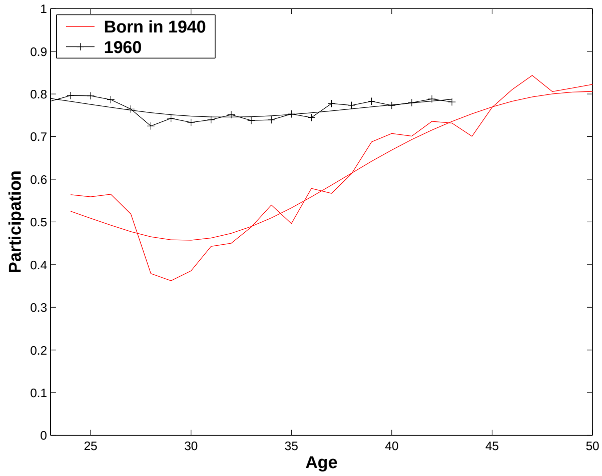
Employment profiles over the life-cycle and their changes across cohorts differ considerably by education. They are mostly increasing in age for high school graduates, while college women tend to work a lot before childbearing ages, then drop out of the labor force and finally join the labor force again after childbearing ages (see Figures 17 and 18, respectively).

Fig. 17: Life-Cycle Employment Profile of Married Women by Cohort - High School Graduates



However, changes in employment rates across cohorts are the largest at childbearing ages. Between age 20 and 35, employment rates increased on average by 26 and 27 percentage points for high school and college graduates respectively. Between age 36 and 50, employment only increases by 10 and 4 percentage points for HS and College graduates.

Fig. 18: Life-Cycle Employment Profile of Married Women by Cohort - College Graduates



7.1.2 Fertility

In Table 7, we present the total number of children ever born and the average age of mother at birth of first child by education. Fertility levels decreases with education. However, changes are the largest for High School graduates women. On the other hand, age of mother at birth of first child increases with education and changes are the largest for College educated women.

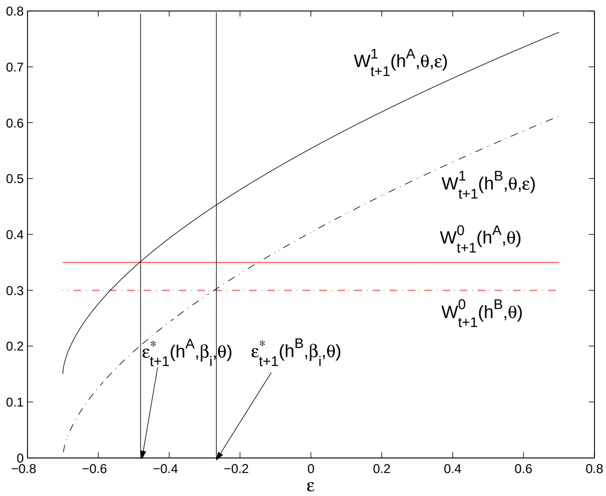
Tab. 7: Fertility Levels and Timing of Births by Education (Std. Dev.)

	1940 Cohort	1960 Cohort
Total Number of Children Ever Born:		
High School	2.6 (1.2)	2.1 (1.1)
College	2.3 (1.2)	1.8 (1.1)
Age of Mother at Birth of First Child:		
High School	23.3 (2.7)	25.5 (4.7)
College	25.9 (3.7)	29.2 (4.6)

7.2 Comparative Statics

We derive the comparative statics of ϵ^* with work experience, h, coefficients of Mincer wage equation, $\{\beta_i\}_{i=0,1,2}$, and women's type, θ. Consider two women of type θ who have the same number of years of work experience in period t, $h_t^A = h_t^B$. Suppose that woman A receives a productivity shock above the reservation shock, while woman B receives a shock below. Next period, we have $h_{t+1}^A = h_{t+1}^B + 1$. This affects their respective continuation values as well as the period utility if they decide to work (since the wage is increasing in the number of years of work experience). The indirect utility both, from working and not working, is higher for woman A than woman B (see Figure 19), but the difference in the former is larger than the difference in the latter. Hence, $\epsilon_{t+1}^*(h^A, \beta_i, \theta) < \epsilon_{t+1}^*(h^B, \beta_i, \theta)$ and woman A is more likely to work in the future. By analogy, the same comparative statics apply for β_i, holding everything else the same. Women are more likely to work, following a decrease in the gender wage gap due to an increase in β_0 or an increase in the returns to experience due to an increase in (β_1, β_2).

Fig. 19: Comparative Statics - $h^A > h^B$



Next, consider two women who have a different number of children over their lifetime, $f^A < f^B$, but the same age at birth of first child, $a^A = a^B$. Here comparative statics depend on the relative magnitudes of time and goods costs. For sake of intuition, consider two extreme cases. In the first case, children are costly if the woman works, while in the second, they are costly if she doesn't work. Therefore, in case (1) (case (2)), women with fewer children are more (less) likely to work. In the data section above, we described employment by number of children in the household and found that it is decreasing. Calibrating to this fact, parameters adjust such that a version of case (1) applies. Therefore, under our parameters, the reservation shock is higher for women who have more children over their life-time. Thus, they are less likely to work. Finally, consider two women who differ in their age at birth of first child, i.e. $a^A > a^B$, but will have the same number of children over their life-time, $f^A = f^B$. Then for some periods, woman A will have fewer children than woman B. She will therefore be more likely to work, everything else the same. However, once she starts to have children herself, she will have younger children in the household than woman B. Since younger children are costlier, she is less likely to work during those periods.

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