# To Grandmother's House We Go: Childcare Time Transfers and Female Labor Mobility*

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**Abstract**

Women in the United States frequently rely on childcare from extended family but can only do so if they live in the same location as them. This paper studies how child care costs, the location of extended family, and fertility events influence both the labor force attachment and labor mobility of women in the United States. We begin by empirically documenting strong patterns of women returning to their home locations in anticipation of fertility events, indicating that the desire for intergenerational time transfers is an important motivator of home migration. Moreover, women who reside in their parent's location experience a substantial long-run reduction in their child earnings penalty. Next, we build a dynamic model of labor force participation and migration to assess the incidence of counterfactual scenarios and childcare policies. We find that childcare subsidies increase lifetime earnings and labor mobility for women, with particularly strong effects for women who are ever single mothers and Blacks. Ignoring migration can understate the welfare benefits of these policies by a meaningful extent.

**Keywords:** Migration, childcare, female labor supply, human capital.

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# 1 Introduction

How does childcare availability influence the labor force attachment and migration behavior of women in the United States? The cost of childcare in the United States is widely acknowledged to be a non-trivial financial hardship for many families: recent surveys indicate that the average cost of center-based infant care exceeds 27 percent of median income for single parents (Child Care Aware, 2017). One option that families may choose in response to these high costs is living near their own parents to use take advantage of cheap or free child-care — informal care is commonly used in the U.S.[1] but can only be used if relatives are nearby. Thus, childcare needs may constrain both the labor force *participation* and the labor *mobility* of U.S. women.

The central goal of this paper is study how migration choices are constrained by childcare needs and the implications for these location constraints for women's earnings. We begin with an empirical analysis that documents strong patterns of women moving to their birth state in anticipation of fertility events, suggesting that childcare assistance indeed plays a role in motivating home migration. Migratory mothers who are not moving back to their home location additionally appear to prefer states with lower child care costs. We also find that women with children exhibit considerably stronger labor force attachment when living in their home state or in states with lower child care costs. Lastly, we use panel data from the Panel Study of Income Dynamics to explore the impacts of a birth on women's earnings using an event study style design (a la Kleven et al., 2019) and analyze how this 'child penalty' varies based on the mother's proximity to her own mother and on the local child care costs. We show that women who give birth in the same state as their own parents experience a substantially smaller child earnings penalty than women who have a child elsewhere, suggesting that these immediate responses have long-run implications. In all these cases, the effects are stronger among unmarried women.

These descriptive facts motivate the question of how subsidized child care would alter the migration and working decisions of women in the U.S. Typically, analyses of the impacts of child care subsidies focus on the direct impacts of such subsidies on labor force participation and human capital accumulation as the primary mechanism through which such subsides impact women's earnings.[2] However, if the high costs of non-subsidized child care prevent households from moving far from their parents and thus from optimally sorting across labor

[1] Roughly 20% of families with young children report use relative-provided child care McMurry (2021).

[2] For an overview of the literature on the elasticity of women's labor supply to child care costs, see (Del Boca, 2015).

markets, we might expect that there is a secondary effect on earnings and welfare stemming from reduced frictions in labor mobility.

To explore this mechanism, we construct and estimate a model that nests a canonical model of dynamic labor force participation in a model of dynamic migration. Women receive shocks to their fertility and marriage status at the beginning of every period, after which they choose their labor force attachment and whether/where to move. The women in the model with children must balance the trade-off between building experience (Eckstein and Wolpin, 1989) through labor force participation and paying more in child care costs, though they retain the option to move to their parent's location so as to receive their assistance. Results from the model suggest that fully subsidizing childcare would increase lifetime wages for women who are ever single parents by over 10 percent on average and over 5.8 percent for women who are never single parents. These subsidies also encourage labor mobility for single mothers substantially, raising lifetime labor mobility by 3 percent. Ignoring migration when estimating the welfare gains of the policies results in understating them by a considerable extent, particularly for single mothers. Moreover, we find that labor mobility increases in counterfactual settings where children are born only to married women, suggesting that the recent increases in the share of single-parent families may have played a role in concurrent declines in labor mobility. We also estimate our model separately for Black and white women and find considerably stronger effects for the former group.

Our paper expands upon and ties together three broad areas of research: the literature on child care costs and women's labor force participation, the literature on the determinants of migration, and the literature on the implications of family-based ties for labor market outcomes.

First, our paper introduces a new mechanism that contributes to the 'child penalty' faced by mothers: increased job mobility frictions caused by location-specific child care access. A long existing literature has documented the fact that women experience large dips in earnings following the birth of a child (Kleven et al., 2019, Cortes and Pan, 2020, Goldin and Mitchell, 2017, Budig and England, 2001, Angrist et al., 1998). One explanation for mother's dip in earnings is that the lack of affordable child care forces women either out of the labor force or into part-time work, resulting in periods of low or zero earnings and lower earnings growth over time to due slower human capital accumulation. Analyses of programs which provide free or subsidized child-care/early childhood education in Canada (Baker et al., 2008, Lefebvre and Merrigan, 2008), Europe (Bauernschuster and Schlotter, 2015, Bettendorf et al., 2015, Havnes and Mogstad, 2011, Lundin et al., 2008), and the United

States (Cascio, 2009, Tekin, 2007, Blau and Tekin, 2007, Borowsky et al., 2022) indicate that subsidized child care increase the likelihood that women work, while also crowding out their use of informal care. This crowd-out is usually discussed in context of changing the quality of care received by children or as reducing the elasticity of women's labor supply with respect to care subsidies. We argue, however, that this substitution away from informal care is also a mechanism through which these subsidies may improve women's labor market prospects. By allowing women to no longer rely on relative care, they are able to be more mobile and potentially achieve welfare gains by moving to a more productive labor market than their parents live in.

By incorporating this mechanism into a dynamic model of labor force participation and fertility, we are building upon a long-standing strand of the women's labor participation literature which considers how child care would change women's labor force participation decisions throughout the life-cycle rather than just in the immediate aftermath of policy implementation. Our model builds directly on the frameworks of dynamic labor supply in presence of fertility seen in Eckstein and Wolpin (1989), Francesconi (2002), Haan and Wrohlich (2011), and Bick (2016). The latter two do incorporate private child care costs, but we extend these models by incorporating informal care from spouses and grandparents as well as adding in a migration component. By incorporating these components, we aim to consider how child care subsidies may change women's welfare not only through increasing their labor force participation, but by changing the labor markets in which they supply labor.

This approach is complementary to the analysis seen in Adda et al. (2017) which considers how women make decisions about both labor force participation and the occupation in which to work. Their model suggests that while three-quarters of the career costs of children are attributable to reduced labor supply, part of the loss is attributable to occupation choices in anticipation of fertility that reduce earnings. Our model considers a different margin — location, rather than occupation — through which women may be adjusting their labor supply to account for child care needs. Our results similarly suggest that while the majority of the welfare gains women might accrue from child care subsidies are from increased participation, a meaningful proportion of them may also be attributable to women being able to sort into their preferred labor markets.

Second, our model also contributes to our understanding of the factors influencing return migration and home-biases in location choices. Older work has studied repeated and return migration (Davanzo, 1983; Dierx, 1988) with the view that such moves are driven entirely by monetary influences. Some more recent work (Diamond, 2016; Kennan and Walker, 2011;

Bishop, 2008) considers non-monetary factors agents weigh when making repeated moving and location choices, but these papers typically condense preferences for living in one's home location into a single utility premium. A small literature has documented the role of emotional attachment to places' characteristics and the role of concentration of extended family in location decisions. (Boyd et al., 2005; Spilimbergo and Ubeda, 2004; Zabek, 2019; Spring et al., 2017). Through focusing on fertility as a new driver of home migration, we aim to contribute to the endeavor to unpack the specific determinants of return migration and add to the literature that studies how individuals balance pecuniary and non-pecuniary factors when making migration decisions in the United States.

In particular, these analyses may help to understand the factors underpinning recent declines in long-distance migration (Molloy et al., 2011). Recent research (Johnson and Schulhofer-Wohl, 2019) suggests that declines in the long-distance migration rate in recent generations is primarily a consequence of a decline in return migration. Johnson and Schulhofer-Wohl focus, however, on a different definition of return migration than the current paper – a move to any location one once lived in, rather than a move to the location one was raised in. Nonetheless, our results point to recent declines in U.S. fertility rates as a potential component of this observed drop in return migration.

Lastly, by focusing on home-based return migration, our results also marry the literature on migration with a growing literature on the implications of family-based ties for labor market outcomes. Proximity to family can mitigate child or elder care needs, allowing greater attachment to the labor force. Geographic distance from one's mother or mother-in-law is associated with a greater likelihood of child care transfers, allowing for higher labor force participation for women (Compton and Pollak, 2015, 2014; Chan and Ermisch, 2015). To identify the effects of access to grandparent care, past research has used variation in pension generosity and retirement age (Dimova and Wolff, 2011; Aparicio-Fenoll and Vidal-Fernandez, 2015; Zamarro, 2020; Bratti et al., 2018; Posadas and Vidal-Fernandez, 2013) and the death of grandparents (Arpino et al., 2014; McMurry, 2021) to show that larger grandparent time transfers are associated with higher labor force participation and earnings for mothers. Beyond the realm of child care, co-location near parents acts as a buffer against earnings losses for adult children following a job displacement (Coate et al., 2017; Kaplan, 2012). Conversely, care needs may flow in the opposite direction, with adult children living near parents having greater care responsibilities for aging or ill parents, resulting in worse economic outcomes (Charles and Sevak, 2005; Konrad et al., 2002; Rainer and Siedler, 2009).

To our knowledge, the only other paper that directly assess the role of informal child care in influencing the migration choices of women is García-Morán and Kuehn (2017), who build a model of residence choice, fertility decisions, and female labor force participation in the context of Germany. Our contribution relative to their paper comes from our focus on dynamics: the authors model migration, working, and fertility decisions as one-shot choices the agent solves at the start of the model. However, labor force participation and migration are dynamic processes[3], which means our dynamic framework will better capture the life-cycle implications of childcare availability and policies. Our model of migration decisions is most complementary to Coate (2013), who considers a dynamic model of migration where agents take the location of their parents into account and are willing to accept lower wages in exchange for closer proximity to their parents. His model, however, focuses on early adulthood migration decisions and does not take fertility into consideration.

The paper is organized as follows: Section 2 motivates our research question by providing descriptive evidence regarding the timing of home migration and fertility events observed in U.S. data. Section 3 details our model, and Section 4 describes our estimation procedure. Section 5 presents model estimates and evaluates the model's fit, while Section 6 presents the results of counterfactual simulations. Finally, Section 7 considers potential avenues for future research before concluding.

# 2 Motivation

In this section, we present empirical evidence to suggest that U.S. women respond to the incentives discussed in the introduction. We begin by showing in the American Community Survey (Ruggles et al., 2020) that U.S women frequently return to the birth state (which we take as a proxy for their parent's location for lack of a better alternative) in anticipation of fertility events and that those who have children in their home state exhibit markedly higher labor force attachment than those who live elsewhere. Next, to further motivate our focus on dynamics, we construct event-study representations of the child earnings penalty in the style of Kleven et al. (2019) using the Panel Study of Income Dynamics (PSID) and show that women who live in the same state as their grandparents experience a considerably smaller long-run child penalty than those who do not.

[3] Multiple moves and return migration are salient features of the data (Kennan and Walker, 2011).

## 2.1 Fertility and Return Migration among U.S. Women

With how costly child care is in the United States, one may expect women with small children to make different location and working decisions than those without. In particular, we may expect women with young children to be more likely to move back to their parent's location to take advantage of familial support in raising young and for women with children to work more hours if their parents are in their same location than if not. Women with small children should also be more reluctant to move to locations with higher childcare costs than those without, other things equal.

To test these hypotheses, we use data from the 2005-2017 waves of the American Community Survey (ACS). Each year of the ACS contains a 1 percent sample of the entire United States' population, providing a large number of observations. Additionally, the 2005-onward waves of the ACS also contain information on one-year migration histories (or, the state that respondents were in the year before their interview). While somewhat limited, this information, coupled with extensive demographic information and state-level measures of child care costs, will allow us to observe some simple margins of behavior that the presence of young children influences.

We restrict our ACS sample to women aged 22-35 who were born in the United States. We drop individuals that did not complete at least one year of high school education and also exclude observations that either report working more than 75 hours per week on average or who have negative income. The women in our sample are limited to those who are coded as household heads, spouses of household heads, or children/children-in-laws of household heads (to allow for the possibility of "boomerang migration," or individuals moving back into their parents' home). The ACS additional records the youngest own child in for all respondents, allowing us to distinguish women who have young children from those who do not. We exclude observations whose age and age of youngest child imply a birth before the respondent was age 14.

We first investigate whether women are more likely to move home in response to fertility events. We restrict our sample to women who were not living in their state of birth in the year before the interview and then run the linear probability model:

$$h_{it} = \beta_0 + \boldsymbol{\beta}_1 \boldsymbol{X}_{it} + \beta_3 f_{it} + \tau_t + \varepsilon_{it},$$

where $h_{it}$ indicates whether individual $i$ moved back to their birth state in year $t$.[4] $\boldsymbol{X}_{it}$

[4] The variable is scaled to be either 0 or 100 — thus, regression coefficients can be interpreted simply as

Table 1: Effects of First Pregnancy on Home Migration Probability (HMP)

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| VARIABLES | HMP | HMP | HMP | HMP |
| Mean Dep Var | 4.03 | 4.44 | 3.58 | 4.44 |
| First Pregnancy (FP) | 0.608 | 2.237 | 0.259 | 2.076 |
| | (0.198) | (0.525) | (0.204) | (0.687) |
| FP × High Childcare Costs | | | | 0.361 |
| | | | | (1.059) |
| Age | -1.573 | -1.817 | -1.749 | -1.816 |
| | (0.141) | (0.194) | (0.233) | (0.194) |
| High School Degree | 0.0123 | -0.0538 | 0.0639 | -0.0530 |
| | (0.195) | (0.253) | (0.304) | (0.253) |
| College Degree | 0.863 | 1.525 | 0.219 | 1.524 |
| | (0.194) | (0.255) | (0.302) | (0.255) |
| Sample | All | Non-Married | Married | Non-Married |
| Observations | 572,964 | 279,471 | 293,493 | 279,471 |
| R-squared | 0.008 | 0.010 | 0.008 | 0.010 |

**Notes:** Robust standard errors in parentheses. Sample is US-native women aged 22-35 in the 2005-2017 ACS who completed at least one year of high school and were not located in birth state the previous year. Additional controls include fixed effects for birth state and calendar year, a quadratic in age, an indicator for some college attained, amenity measures for state lived in last year (college share, unemployment rate, rates of violent and property crime, population, per-capital government student expenditure, student-teacher ratios, and share of days warmer than 70 degrees) and Black and hispanic indicators. First pregnancy indicator defined by presence of a child less than one year old while being the only own child of the respondent in the household. Regressions weighted by sampling weights.

contains a vector of demographic controls, while $f_{it}$ indicates individual $i$'s first fertility status in year $t$, defined by presence of a child belonging to the respondent that is less than 1 year old while also being the only child of the respondent in the household. The term $\tau_t$ contains year fixed effects, while $\varepsilon_{it}$ is an error term. Standard errors are corrected for heteroskedasticity, and regressions are weighted using sampling weights provided by the ACS. We focus on the first pregnancy because the presence of additional children may make migration more cumbersome — thus, women may be more likely to move home in response to their first fertility event than subsequent ones. We also run our specification for all women as well as for non-married and married women separately, as the additional spousal financial support available to married women may make them less likely to migrate in response to

percentage point changes to the likelihood of a home move.

Table 2: Effects of Young Child on Home Migration Probability (HMP)

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| VARIABLES | HMP | HMP | HMP | HMP |
| Mean Dep Var | 4.03 | 4.44 | 3.58 | 4.44 |
| Young Child (YC) | -0.0212 | 0.174 | 0.0240 | 0.156 |
| | (0.0743) | (0.134) | (0.0889) | (0.175) |
| YC × High Childcare Costs | | | | 0.0398 |
| | | | | (0.260) |
| Age | -1.570 | -1.844 | -1.752 | -1.843 |
| | (0.141) | (0.195) | (0.233) | (0.195) |
| High School Degree | 0.0132 | -0.0259 | 0.0676 | -0.0260 |
| | (0.195) | (0.253) | (0.304) | (0.253) |
| College Degree | 0.867 | 1.558 | 0.234 | 1.556 |
| | (0.194) | (0.258) | (0.301) | (0.258) |
| Sample | All | Non-Married | Married | Non-Married |
| Observations | 572,964 | 279,471 | 293,493 | 279,471 |
| R-squared | 0.008 | 0.010 | 0.008 | 0.010 |

**Notes:** Robust standard errors in parentheses. Sample is US-native women aged 22-35 in the 2005-2017 ACS who completed at least one year of high school and were not located in birth state the previous year. Additional controls include fixed effects for birth state and calendar year, a quadratic in age, an indicator for some college attained, amenity measures for state lived in last year (college share, unemployment rate, rates of violent and property crime, population, per-capital government student expenditure, student-teacher ratios, and share of days warmer than 70 degrees), and Black and hispanic indicators. Young child defined as presence of own child aged at most 4 in household. Regressions weighted by sampling weights.

fertility than single women.

Tables 1 and 2 report the results of this exercise for two differen specifications of $f_{it}$. In Table 1, $f_{it}$ is a dummy variable equal to 1 if the respondent had their first child in the previous year, which may be observed by the presence of a child of the respondent's that is less than 1 year old while also being the only child of the respondent in the household. In Table 2, $f_{it}$ is a dummy variable equal to 1 if the respondent has any children four years old or younger in the household. Conceptually, we may the presence of children to make migration more cumbersome and costly, so women may be more likely to move home in response to their first fertility event than subsequent ones. At the same time, the additional spousal financial support available to married women may make them less likely to migrate in response to fertility than single women.

These predictions are well born-out in the data. While Table 2 indicates that the presence of small children does not meaningfully influence the likelihood of a home move, Table 1 suggests that initial fertility events make women noticeably more likely to home-migrate. These effects are also much stronger for single women than married women — indeed, the subgroup analysis indicates that married women respond to initial fertility events by migrating quite little. However, with a base home rate of migration of around 4 percent in the data, initial fertility events make single women roughly 50% more likely to move home compared to the rest of the sample. We also find that having previously lived in a state with high child care costs[5] is associated with a slightly higher likelihood of moving back home, though this effect is not statistically significant.

While higher child care costs do not seem to substantially influence the extensive margin of probability of a move itself, they may still have intensive margin impacts in that they could distort the location choices of women conditional on moving in the first place. We next investigate whether the presence of young children distort the location choices of women who choose to migrate. We limit our ACS sample to women who are observed to have moved from their previous-year state and have not moved to their state of birth. We then test whether the presence of young children result in women being less likely to locate in high childcare-cost states, defined as being above-median as before. Table 3 presents the results of this test and affirms the hypothesis — moving women with young children on average choose to locate to states with lower childcare costs than those without, with the effects again being noticeably stronger for single women than married ones.

Finally, we investigate how location and the presence of children influence the labor force attachment of women in the ACS. The ACS records usual hours worked per week for all employed respondents — for unemployed respondents or respondents not in the labor force, we code usual hours worked per week as zero. We then regress usual hours worked per week on a variety of covariates to do with the presence of children, location, child care costs, and marital status. Intuitively, higher child care costs ought to decrease hours worked by women because it makes working relatively more expensive. Being proximal to parents ought to increase labor force attachment if parents primarily provide time transfers in child-rearing, but effects of marital status on labor force attachment are a-priori ambiguous. Women with husbands may exhibit higher labor force attachment if their husbands also provide time

[5] Defined as having above median costs, with numbers coming from Child Care Aware (2017), who survey state child care resource and referral networks to obtain average prices for full time child care centers for three age groups in each U.S. state. For a visual representation of average full-time infant childcare expenses across U.S. states, refer to Figure A.1a.

Table 3: Effects of Children on Probability of Moving to High CCC State (HCS)

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| VARIABLES | HCS (0-100) | HCS (0-100) | HCS (0-100) | HCS (0-100) | HCS (0-100) |
| Mean Dep Var | 50.48 | 52.86 | 47.96 | 42.86 | 47.96 |
| Young Child | -1.693 | -4.816 | -0.412 | -4.647 | -0.387 |
| | (0.654) | (1.271) | (0.763) | (1.258) | (0.760) |
| High School Degree | 3.895 | 2.590 | 5.240 | 2.153 | 5.519 |
| | (1.897) | (2.521) | (2.831) | (2.518) | (2.796) |
| College Degree | 13.52 | 10.88 | 15.07 | 10.61 | 15.28 |
| | (1.853) | (2.478) | (2.786) | (2.481) | (2.751) |
| Sample | All | Non-Married | Married | Non-Married | Married |
| Home State FE | YES | YES | YES | YES | YES |
| Previous State FE | NO | NO | NO | YES | YES |
| Observations | 54,233 | 26,582 | 27,651 | 26,582 | 27,651 |
| R-squared | 0.069 | 0.088 | 0.052 | 0.106 | 0.063 |

**Notes:** Robust standard errors in parentheses. Sample is US-native women aged 22-35 in the 2005-2017 ACS who completed at least one year of high school and who moved in the previous year and not to their state of birth. Additional controls include fixed effects for birth state and calendar year, a quadratic in age, an indicator for some college attained, amenity measures for state lived in last year (college share, unemployment rate, rates of violent and property crime, population, per-capital government student expenditure, student-teacher ratios, and share of days warmer than 70 degrees), and Black and hispanic indicators. Young child defined as presence of own child aged at most 4 in household. Regressions weighted by sampling weights.

assistance in raising their progeny, but if a husband's primarily role is to slacken budget constraints by providing supplementary income then we ought to see married women with children work less than single women, other things equal.

Table 4 presents the results of this exercise, with many of the above predictions clearly manifesting in the data. The presence of children decreases usual weekly hours worked by women substantially, and the effects are noticeably stronger for married women. However, women who have children in their birth state work more than women who do not, while women with children in states with higher child care costs also work relatively less.

## 2.2 Grandparent Proximity and the Child Penalty

While our analyses using ACS data provide a snapshot of the impacts of young children on household location choices, cross-sectional data cannot tell us the long-term impacts of living near relatives or living in low child-care cost regions on women's lifetime earnings trajectory. It is well-documented that women experience a decline in earnings following births, often

Table 4: Effects of Children and Location on Hours Worked

| VARIABLES | (1) Hours | (2) Hours | (3) Hours | (4) Hours |
|---|---|---|---|---|
| Mean Dep Var | 27.27 | 27.27 | 27.80 | 26.64 |
| Married | 0.291 | 0.292 | | |
| | (0.0403) | (0.0403) | | |
| Child Present (CP) | -2.640 | -4.436 | -2.902 | -9.411 |
| | (0.0574) | (0.0875) | (0.128) | (0.0960) |
| In Home State | 0.609 | -0.324 | -1.058 | 1.270 |
| | (0.0364) | (0.0430) | (0.0548) | (0.0692) |
| CP × Married | -5.577 | -5.475 | | |
| | (0.0723) | (0.0723) | | |
| CP × High Childcare Costs | | -0.502 | -0.640 | -0.683 |
| | | (0.0665) | (0.112) | (0.0882) |
| CP × In Home State | | 2.726 | 1.036 | 2.260 |
| | | (0.0756) | (0.133) | (0.0996) |
| Sample | All | All | Non-Married | Married |
| Observations | 2,056,614 | 2,056,614 | 1,014,536 | 1,042,078 |
| R-squared | 0.117 | 0.118 | 0.116 | 0.128 |

**Notes:** Robust standard errors in parentheses. Sample is US-native women aged 22-35 in the 2005-2017 ACS who completed at least one year of high school.Additional controls include fixed effects for birth state and calendar year, a quadratic in age, an indicator for some college attained, amenity measures for state lived in last year (college share, unemployment rate, rates of violent and property crime, population, per-capital government student expenditure, student-teacher ratios, and share of days warmer than 70 degrees), and Black and Hispanic indicators. Young child defined as presence of own child aged at most 4 in household. Regressions weighted by sampling weights.

referred to as the 'child penalty,' which persists for up to ten years post birth (Kleven et al., 2019), This child penalty is common across a number of European and North American countries (Kleven et al., 2019), ranging from a 20% decline in women's earnings relative to pre-birth earnings in Scandinavian countries to 44% in the US or 60% of pre-birth earnings in Germanic countries. A large factor in the decline in earnings is women's withdrawal from the labor market. Therefore, we might expect that having access to cheaper or free child care would allow women to work more hours and reduce the child penalty.

To test this, we use data from the Panel Study of Income Dynamics (PSID) to estimate the size of the child penalty for women living near or far from grandparent care and for women living in high vs. low child-care cost regions. We adopt a modified form of the event study specification first proposed by Kleven et al. (2019). Because the authors show that there is no child penalty for men, we focus solely on women's first births. For each mother in the data, we define event time ($t$) based on the year of their first child's birth. Our outcome of interest is woman $i$'s earnings $Y_{ist}$ in year $s$ and at event time $t$. The regression is as follows:

$$Y_{ist} = \sum_{j \neq -1} \alpha_j \mathbb{1}[j = t] + \sum_k \beta_k \mathbb{1}k = age_{is}] + \sum_n \gamma_n \mathbb{1}[n = s] + \epsilon_{ist}. \quad (1)$$

The regression contains event-time dummies with $\alpha$ coefficients, age dummies with $\beta$ coefficients to control for life-cycle trends, and year dummies with $\gamma$ coefficients to control for time trends. Event-time $t = -1$ is omitted, so all estimates are relative to the year just prior to birth. As noted in Kleven et al. (2019), we are able to identify effects of all three sets of dummies because of the variation in the age at which women have children.[6]

The parameters of interest are the $\alpha$ parameters, but they will represent differences in levels. To transform them into percent changes, we calculate $P_t = \frac{\hat{\alpha}_t}{\hat{\mathbb{E}(Y_{ist}|t)}}$, where the bottom of the fraction is the predicted outcome omitting the contribution of the event-time dummies.

We estimate this regression for all women living in the same state as the mother of the mother (henceforth the grandmother), all woman living in different states than the grandmother, women living in states in the top half of the childcare cost distribution, and women living in states in the bottom half of the childcare cost distribution. For these comparisons of the child penalty across groups, our figure of interest is the child penalty gap:

$$\frac{\hat{\alpha}_t^1}{\hat{\mathbb{E}(Y_{ist}^1|t)}} - \frac{\hat{\alpha}_t^2}{\mathbb{E}(\hat{Y}_{ist}^2|t)}.$$

We use the Delta method to calculate standard errors of this gap and then test whether we can reject the null that the child penalty is equal for those living in the same state as the grandmother and those living in different states.

Note that these estimates should not be interpreted as the causal impact of living near a grandparent or in a child care cost region on the child penalty. We expect that women are sorting across these locations in part based on their attachment to the labor force; women who want to continue working after a birth for reasons unobservable to us as econometricians are more likely to settle in places with affordable child care, whether that be relative care or cheaper private care options. We cannot separate these indirect selection effects from the direct effects of having cheaper child care available. Nonetheless, these patterns will provide suggestive evidence of whether child care cost factors are meaningfully related to the

[6]For more details on the identification assumptions needed to assume this is the causal impacts of child birth, see Kleven et al. (2019).

long-term child penalty women face following their first birth.

### 2.2.1 Data

For this analysis, we need a panel of income data for women in the years prior to and the years following their first birth. To create this, we use the PSID's full retrospective history of births and adoptions, which provides the full history of births for those interviewed in the years 1985 onward. Using this data set, we create a sample of all PSID women who have at least one birth, the year their first birth occurred, their age at that birth, and whether they were married at the time of that birth. We then combine this data with information from the PSID family files on earned income in each year of the women's life, the US state they live in in each year, and the US state that their parents live in each year. Earned income is defined as the reported total income including wages and other income.[7]

Following the restrictions used by Kleven et al. (2019), the panel of years includes five years pre-birth and ten years post-birth. Women are excluded from the sample if they are missing more than 8 years in this period, missing all years pre-birth, or all years post-birth. We also restrict the data to be from 1985 onward in part to reduce measurement error from retrospective birth histories and in part to match the data sample cleaned for the estimation sample, which only contains locations from 1985 onward.

We look at differences in the child penalty across three categories of mothers, as well as the interaction between these categories:

1. Near vs. Far to Grandma: A woman is near to Grandma if the mother is in the same state as her own mother in the year of her first birth.

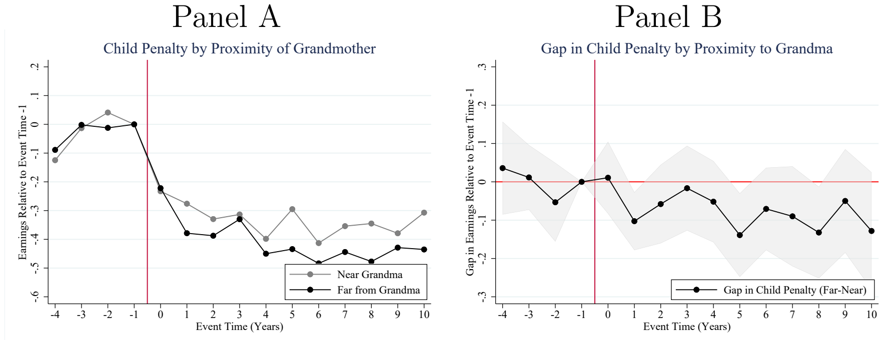
2. High vs. Low Child Care Costs: A woman is a high child care cost type if she lives in a state that has child care costs above the median of our CC cost index in the year of her first birth.

3. Married vs. Unmarried: A woman is married if she was married in the year of her first birth.

[7] For women who were household heads, this is based on reported income from 'wages and other income". For women who were spouses in the data, the measurement of income changes in 1993 when they begin separating out business and farm income. Due to this change, we create total income for spouses by adding together the total income excluding business and farms and total income from businesses for all years post 1993. Income is then assigned by the sex of the head of household.

Figure 1: Child Penalty for Women Living Near or Far from Grandmother

Panel A Panel B



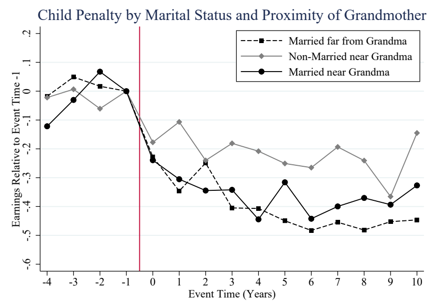

**Notes:** Figure 1A (left) plots coefficients from event studies of earnings on indicators for years surrounding a woman's first birth for both women who live in the same state as the grandmother (near) or different states (far). The unit are percent changes (0 to 1) in earnings relative to the year prior to birth. The regression includes controls for age of mother at first birth and year of birth. Figure 1B (right) calculates the gap for those near vs. far and reports 10% confidence intervals for a test of the null that this gap is equal to zero.

### 2.2.2 Results

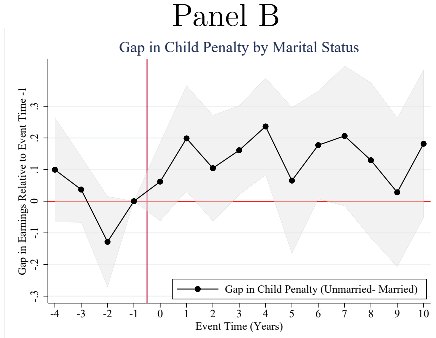
Figure 1 plots the coefficients from the event studies described in (1) with panel A plotting the coefficients separately for mothers living in the same state or in different states from the child's grandmother and panel B plotting the size of the gap between these groups. While both types of mothers experience a large child penalty, those living distant from the grandmother experience a 10 percentage point larger child penalty that persists for up to 10 years following the child's birth. When we split the results by marital status (Figure 2), we see that the effects are driven primarily by smaller earnings losses for single mothers living near the child's grandmother rather than married mothers.

Figure 2: Child Penalty for Women Living Near or Far from Grandmother, by Marital Status

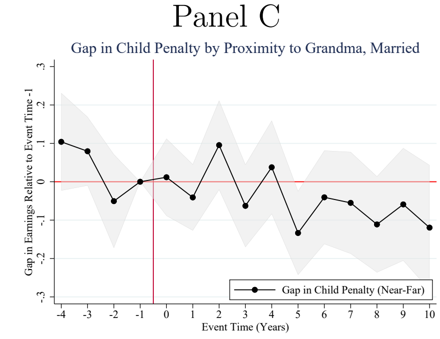
Panel A




Panel B




Panel C




**Notes:** Figure 2A (top) plots coefficients from event studies of earnings on indicators for years surrounding a woman's first birth for married women who live in the same state as the grandmother, married women who live in different states, and unmarried women who live in the same state as the grandmother. The unit are percent changes (0 to 1) in earnings relative to the year prior to birth. The regression includes controls for age of mother at first birth and year of birth. Figure 2B (bottom left) calculates the gap in the percent decline for those near the grandmother who are married vs. unmarried. Figure 2C (bottom right) calculates the gap in the percent decline for those married near vs. far from the grandmother. Both report 10% confidence intervals for a test of the null that this gap is equal to zero.

Table 5 reports the results of a regression which aggregates the coefficients into pre-period (excluding the year prior to birth), year of birth, and post-period and interacts these periods with an indicator for living near the grandmother. Here, we see that the effects of living near the grandmother are statistically significant in the full sample and recoups about \$2,700, or 22% of the child penalty faced by mothers. In contrast, for single mothers the effects are much larger, around \$8,400 or 66% of the child penalty faced by single mothers. The benefit of living near grandmothers is smaller and not statistically significant for married mothers.

Table 5: Aggregated Child Penalty, by Distance to Grandmother

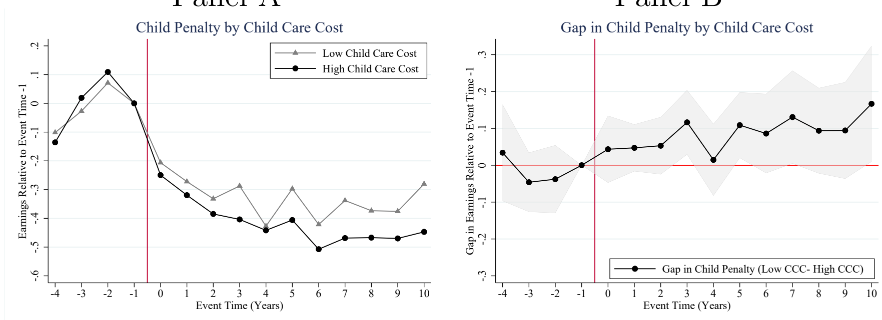
| | (1)<br>Full Sample | (2)<br>Unmarried | (3)<br>Married |
|---|---|---|---|
| Pre-period | -430.3 | -7218.8** | 493.8 |
| | (1096.6) | (2714.4) | (1174.1) |
| Year of birth | -5113.4*** | -5644.1 | -5094.2*** |
| | (1262.3) | (4771.6) | (1279.9) |
| Post-period | -12068.9*** | -12836.3*** | -12454.1*** |
| | (1042.8) | (3124.9) | (1135.1) |
| Near Grandma × Pre-period | 19.38 | 6885.0* | -851.5 |
| | (1331.3) | (2849.9) | (1459.2) |
| Near Grandma × Year of birth | -169.1 | 2541.9 | -662.7 |
| | (1462.7) | (4703.6) | (1572.4) |
| Near Grandma × Post-period | 2721.8* | 8418.1** | 1927.6 |
| | (1203.7) | (3174.8) | (1381.3) |
| Women-Year Obs. | 13530 | 2201 | 11329 |

Note. This table reports the coefficients of a regression of earnings on indicators for years surrounding a woman's first birth, collapsed into the pre-period (2 to 5 years pre-birth), year of the birth, and post-period (1 to 10 years post-birth). The year prior to birth is omitted. All indicators are interacted with an indicator for is the woman is living in the same state as her own mother (Near Grandma). Controls for year of survey and age of mother are also included. Column 1 is the full sample; column 2 are unmarried at year of birth; column 3 are married at year of birth. Standard errors clustered at the individual level in parentheses; * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

We next do a similar exercise for those living in high or low child care cost states. Figure 3 reports the coefficients for event studies of earnings on indicators for years surrounding a woman's first birth for both women who live in the low or high childcare cost states. We see that the child penalty is larger in states with high child care costs. The difference across child care cost regions is of similar magnitude to the difference in the child penalty for those near vs. far from the child's grandmother. Interestingly, the effects of child care on the child penalty seem to primarily occur for married mothers, as shown in Table 6, which reports the aggregated post-birth effects of a child by child care cost region for the full sample (column 1), unmarried mothers (column 2), and married mothers (column 3). While the child penalty

Figure 3: Child Penalty for Women Living in High or Low Child Care Cost States

Panel A Panel B




**Notes:** Figure 3A (left) plots coefficients from event studies of earnings on indicators for years surrounding a woman's first birth for both women who live in the low or high childcare cost states. The unit are percent changes (0 to 1) in earnings relative to the year prior to birth. The regression includes controls for age of mother at first birth and year of birth. Figure 3B (right) calculates the gap in the percent decline for those in high cost relative to those in low cost states and reports 10% confidence intervals for a test of the null that this gap is equal to zero.

is unaffected by child care costs for unmarried mothers, married mothers' child penalty is approximately $4000 larger in a high child care cost state.

# 3 Model

Taken together, these analyses demonstrate the importance of geographic proximity to affordable child care for women's long-run labor market outcomes– whether it be informal care from a grandparent or less expensive private child care. However, in both the analyses, we are not fully accounting for the joint selection process of location, fertility, and labor force participation. For example, when we observe that mothers living in high child care cost regions earn less than those low child care cost regions, it may be that the mothers in low cost regions were motivated to select into those regions due to higher ability or attachment to the labor force that is known to them but unobserved to us as econometricians. Therefore, a model that places some assumptions on the selection process will be required to account for the endogeneity of migration decisions and to evaluate the impact of policy counterfactuals.

Table 6: Aggregated Child Penalty, by Child Care Costs

| | (1) Full Sample | (2) Unmarried | (3) Married |
|---|---|---|---|
| Pre-period | -293.2 | -1731.5 | 427.3 |
| | (676.3) | (999.8) | (774.7) |
| Year of Birth | -4208.6*** | -1194.3 | -4762.2*** |
| | (915.3) | (1984.0) | (1194.6) |
| Post-period | -8435.2*** | -6255.9 | -9075.2*** |
| | (733.6) | (3361.0) | (1158.3) |
| High CCC × Pre-period | -858.9 | 1517.6 | -1564.4 |
| | (1226.2) | (2362.3) | (1324.1) |
| High CCC ×Year of Birth | -1924.4 | -1593.0 | -1465.7 |
| | (1403.0) | (2990.3) | (1659.4) |
| High CCC × Post-period | -3738.6* | 235.0 | -3955.1* |
| | (1514.4) | (4570.2) | (1858.1) |
| $N$ | 9568 | 1674 | 7894 |
| Women-Year Obs. | 9568 | 1674 | 7894 |

Note. This table reports the coefficients of a regression of earnings on indicators for years surrounding a woman's first birth, collapsed into the pre-period (2 to 5 years pre-birth), year of the birth, and post-period (1 to 10 years post-birth). The year prior to birth is omitted. All indicators are interacted with an indicator for is the woman is living in the top half of the state child care cost distribution (High CCC). Controls for year of survey and age of mother are also included. Column 1 is the full sample; column 2 are unmarried at year of birth; column 3 are married at year of birth. Standard errors clustered at the individual level in parentheses; * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

In particular, we are interested in how policies that may substitute for intergenerational time transfers (such as subsidized childcare) would influence the migration decisions and subsequent earnings of women who might otherwise rely on their parents to assist in child-rearing. Using, we will be able to explore the effectiveness of such policies in improving welfare for different types of parents (e.g., single vs. married), as well as decompose any effects on earnings into a direct effect of changes in attachment to the labor force due to child care policies versus the secondary effects of the policies such as allowing households to sort into better paying labor markets.

Lastly, we plan to estimate the model separately by race to explore heterogeneity in the value of these policies for Black mothers relative to White mothers. The frictions associated with child care access may be particularly important in explaining racial gaps in migration rates and wages, as single motherhood is more common for Black mothers. Our reduced form analysis suggests that single mothers are more dependent on geographic proximity of family for access to care. The model will allow us precisely quantify the extent to which fertility events drive migration across demographic groups in the United States and speak to the extent to which recent changes in family structure in the U.S. may be related to ongoing changes in labor mobility.

## 3.1 Setup and Timing of Decisions

Our model adapts the dynamic migration labor force participation of Eckstein and Wolpin (1989) and nests it in a simple framework of dynamic migration (Kennan and Walker, 2011) while incorporating multiple dimensions of family structure. The model is a dynamic discrete choice model that follows the labor force participation and migration decisions of women. We focus on women due to their stronger geographical attachment to their children compared to men and due to the wealth of evidence that points to fertility events being more influential on female labor force attachment than male.

A period is one year. Agents enter the model at age 22 and are at risk of pregnancy until age 35. Between ages 35 to 40, though agents cannot get pregnant, they may either have young children or have no children. After age 35, we additionally assume that the agent's current marital status remains fixed for the rest of the lifecycle. Agents choose whether to supply labor and, afterward, whether and where to move until making a final labor force decision at age 65, after which they accrue no further utility[8]. We select age 22 as the starting point to allow the bulk of higher education choices to be made while pre-empting the prime fertility years of U.S. women[9].

At the beginning of each period, the women in our model observe the location of their parents and stochastic realizations of marriage and fertility[10]. The women then choose whether

[8]So, the final migration decision is made at age 64.

[9]In our estimation sample, 75% of individuals had the same educational attainment at age 35 as they did at age 22. Moreover, while this choice prevents us from being able to account for teenage pregnancies, we observe that in 2006, 90% of first-time mothers were aged at least 23 (Authors' calculations, 2006 ACS).

[10]Fully endogenizing marriage and fertility in the model would be impossibly complicated. We make this assumption to allow us to focus on the influence of childcare availability on female labor force participation and mobility. We assess the robustness of our policy counterfactuals to allowing a simple fertility elasticity effect later in the paper.

to participate in the labor force[11], weighing increased utility from consumption[12] should they choose to vs. preferences for leisure and savings on childcare expenditures should they not. Participation also increases future expected earnings through accumulating work experience. The women then choose whether or where to move — in particular, their migration options include staying in their current location, moving to the state of their parents, or moving anywhere else, which we subsume into the nine Census divisions[13]. Following their migration decision, women enter the subsequent period.

## 3.2 State Variables and Value Functions

Table 7 presents a complete summary of state variables and notation in the model, which are described in more detail in this section. Locations are indexed by $\ell$, with and $\ell^P$ denoting an agent's parent location. The other locations represent the nine Census divisions, each of which are vary by childcare costs $\delta^\ell$, wage effects $\eta^\ell$, and cost of living $\kappa^\ell$. One's college attainment is indexed by $e \in \{0, 1\}$, years of experience by $x$, and age by $a$. College attainment here is assumed to be static, but experience will be allowed to grow endogenously over time.

We now turn to describing notation for family structure. Marital status is denoted by $m \in \{0, 1\}$ and is assumed to evolve entirely stochastically, depending on other elements of the state space. Men make no decisions in our framework and are assumed to inelastically provide monetary and childcare time transfers to their wives. The variable $a_c$ captures the age of the youngest child in the household, provided that they are less than 5 years old.[14]

[11] We focus on participation vs. non-participation both for simplicity and because we observe in the ACS that the share of women working part-time appears invariant to the presence of small children in the household, suggesting that outright participation is the relevant behavioral margin to consider.

[12] We scale consumption so that one unit corresponds to \$2,000, following from full-time work involving working 40 hours per week, 52 weeks per year. This normalization can also be thought of as normalizing the agent to have a single hour of time.

[13] See Appendix C for division definitions. Agents are allowed to live in the same division of their parents while not being in the parent location, as well, so that they do not receive parental child care time transfers. Gemici (2011) uses the same geographic structure. 72% of cross-state moves observed in the data involve cross-divisional moves as well, so our framework allows us to capture the bulk of labor mobility activity. We have estimated a version of the model that instead uses the 48 mainland U.S. states as geography — in addition to dramatically increased computational load, the behavior of the likelihood is somewhat more erratic in trying to rationalize very low-probability moves. However, the main counterfactual results are virtually unchanged.

[14] We currently do not keep track of the number of young children and instead focus on the presence of any at all. Rosenzweig and Wolpin (1980) study the effects of twins on labor force participation and find that women with twins exhibit a labor force participation rate 0.371 pp lower than women without. While these effects are significant, we view their magnitude as small enough to permit the omission for the time

Table 7: Model Notation

| Description | | Values |
|---|---|---|
| Locations | $\ell$ | $\ell^P, \{1, ..., 9\}$ |
| Location Daycare Cost Types | $\delta$ | $\delta^\ell, \ell \in \{1, ..., 9\}$ |
| Location Wage Effects | $\eta$ | $\eta^\ell, \ell \in \{1, ..., 9\}$ |
| Location Costs of Living | $\kappa$ | $\kappa^\ell, \ell \in \{1, ..., 9\}$ |
| College Attainment | $e$ | 0,1 |
| Age | $a$ | [22,65] |
| Age of Youngest Child | $a_c$ | $\emptyset$, [0,4] |
| Years of Experience | $x$ | [0,40] |
| Marital Status | $m$ | 0,1 |
| Fertility Status | $f$ | 0,1 |
| Spouse Wage FE | $\mu_S$ | $\mu_S^L, \mu_S^H$ |
| Previous LFP Status | $p$ | 0,1 |
| Hours | $h$ | 0,1 |
| Time Transfers from Spouse/Parents | $\tau$ | $\tau^S$, $\tau^{P,m}$ |

**Notes:** Table presents model notation. Description of variables and symbolic representations are contained in the first two columns of the table, while the potential values the variables can take are presented in the third column.

The state $a_c = \emptyset$ stands for when the household has no children aged 5 or younger.

Meanwhile, the variable $f$ captures the fertility status of the woman: $f = 1$ indicates pregnancy, i.e., if $f = 1$ in year $t$ then $a_c = 0$ in year $t + 1$ with certainty. Having pregnancy be a known state allows our women to make migration and and labor force participation decisions *in anticipation* of fertility events. Conception, meanwhile, is entirely stochastic and depends on the other elements of the state space.[15] Women are allowed to have multiple children in that their $f$ state may equal 1 even if the household currently contains a young child, in which case $a_c$ will be reset to zero in the subsequent period. We shut down fertility events at age 35, meaning that when women leave the model at age 40 all children have aged out of early childhood. For more details on how the stochastic processes that govern

being. This almost certainly means that we are understating the costs of childcare and the potential effects of subsidies to them in terms of labor force participation and wages.

[15] In 2008, 54% of births among unmarried women aged 20–29 were unintended, compared with 31% of births to married women in the age group (Zolna and Lindberg, 2012). Moreover, even when planned the timing is not always in the control of women; in PRAMS surveys of women who gave birth, about 18% reported that they would have preferred to have had the birth sooner (Maddow-Zimet and Kost, 2020). Additionally, a preponderance of women cite non-economic reasons as the drivers of the choice to conceive (Edin and Kefalas, 2011).

marriage and fertility realizations are determined, refer to Section 4.

Women are assumed to be endowed with a single unit of time period and may choose to work full time ($h = 1$) or not at all ($h = 0$). Spouses are also assumed to be endowed with a fixed effect $\mu_S$ that affects their earnings potential. Subsuming all the state variables outside of the agent's current location into the vector $\Omega$, the value function for a woman without young children in the model is as follows:

$$V(\Omega, \ell) = \max_h \Big\{ \alpha_1(c) + (1-h)(\alpha_2 + \alpha_e e + \alpha_x + \alpha_c c) + \alpha_3 \mathbb{1}(h \neq p) + \alpha_4 \mathbb{1}(\ell = \ell^P) + \boldsymbol{\alpha_\Gamma \Gamma} + \mathbb{E}_{\zeta_{\ell'}}[V'(\Omega, \ell; h)] \Big\}; \quad (2)$$

$$\kappa^\ell c = w_S \mathbb{1}(m = 1) + wh.$$

Thus, $\alpha_1$ rescales utility over consumption in dollars to util terms, and $\alpha_2$ represents a preference for leisure. Preferences for leisure are further modified based on experience ($\alpha_x$) or if the agent has a college degree ($\alpha_e$), and $\alpha_c$ represents a consumption-leisure complementarity that makes married women less likely to work. The parameter $\alpha_3$ constitutes a penalty borne from changing one's labor force participation status (i.e., $p = 0$ and $h = 1$, or $p = 1$ and $h = 0$), allowing the model to account for frictions individuals face in moving in and out of the labor force. Utility premia for currently being in one's parent's location is captured by $\alpha_4$. Locations differ in amenities $\boldsymbol{\Gamma}$ that include average distance to shore (taken from Lee and Lin (2017)), average number of warm-weather days in a calendar year (taken from Kennan and Walker (2011)), and an index of other amenities related to government provisions and quality of life taken from Diamond (2016).

Consumption here is given by the wages of the woman's spouse (assumed to be supplied inelastically and equal to zero if the woman is unmarried) and the earnings of the woman herself. Log wages of the woman and her spouse are given by the following equations:

$$\log(w_S) = \beta_{S,0} + \boldsymbol{\beta}_S \boldsymbol{X}_S + \mu_S + \eta^\ell;$$

$$\log(w) = \beta_0 + \boldsymbol{\beta X} + \eta^\ell + \varepsilon + \xi;$$

$$\varepsilon \sim N(0, \sigma_\varepsilon) \text{ i.i.d}; \quad \xi \sim N(0, \sigma_\xi) \text{ i.i.d.}$$

The vector of observables of the spouse $\boldsymbol{X}_S$ contain a college dummy and a quadratic in experience, while the agent's observables $\boldsymbol{X}$ contain the same standard Mincerian combination along with dummies for having a child aged 0-1 or a child aged 2-4.[16] With the assumption that husbands supply labor inelastically, the terms of the husband's wage equations can be uncovered directly from data if we assume husbands to be identical to their wives in age and schooling level. Meanwhile, the component's of the woman's wage process will be parameters to be estimated. Location fixed effects, $\eta$ are also assumed to be constant across time and equal for men and women, which with the assumption of exogenous male labor supply will allow us to estimate values for $\eta$ outside the model using male wages. Wages offers for women are additionally shocked by a transient component $\varepsilon$ that will be the key factor in determining whether a woman works in a given state and are measured with error $\xi$ assumed uncorrelated with $\varepsilon$.

The final term of (2), $\mathbb{E}_{\zeta_{\ell'}}[V'(\Omega, \ell; h)]$, represents the expected continuation value given the woman's labor force participation decision. Following her choice of $h$, the woman receives a series of location preference shocks that will determine whether and where she moves:

$$V'(\Omega, \ell; h) = \max_{\ell'} \left\{ \beta \sum_{\Omega'} \mathbb{E}_{\varepsilon}[V(\Omega', \ell')]\mathbf{Pr}(\Omega' | \Omega, h, \ell') - \Delta(\Omega, \ell')\mathbb{1}\{\ell' \neq \ell\} + \zeta_{\ell'} \right\}.$$

The agent takes into account possible state state transitions $\Omega'$ and expected next-period utility after solving her optimal hours decision problem and optimizes their choice of next-period location following a series of location preference shocks $\zeta_{\ell'}$ distributed Type 1 Extreme Value with location 0 and the scale parameter normalized to 1. Fertility and marriage transitions are governed by stochastic functions that we calibrate directly from the data. We assume that the woman can no longer become pregnant at age 35 and that their marriage state at age 35 carries on for the remainder of the life cycle. The agent's next value of $p$ (past-period labor force participation) depends on her selection of $h$. The agent's experience $x$ increments by 1 should they choose to work and 0 if they do not, and the agent's age $a$ increments by 1 with certainty. Next-period utility is discounted by the factor $\beta$.

The parameter $\Delta(\Omega, \ell')$ captures moving costs that the agent faces should they have chosen to do so, which itself depends on other elements of the state space. If a woman moves across locations in a period, she must incur moving costs given by

$$\Delta(\Omega, \ell') = \gamma_0 + \gamma_1 e + \gamma_2 \mathbb{1}\{a_c \neq \emptyset\} + \gamma_3 m + \gamma_4 N^{\ell'}.$$

[16] While we abstract away from an explicit part-time choice, this allows the model to be consistent with women potentially preferring more flexible and possibly lower-paying jobs when parenting a small child.

Moving costs involve a fixed cost, are potentialy smaller for college graduates, and are assumed to be larger for married agents and for agents that already have young children. Furthermore, we allow for moves to larger locations ($N^\ell$ represents the population of division $\ell$ in tens of millions) to less costly as in Kennan and Walker (2011).

Finally, a woman with young children enjoys utility:

$$V(\Omega, \ell) = \max_h \Big\{ \alpha_5(c) + (1-h)(\alpha_6 + \alpha_e e + \alpha_x x + \alpha_c c) + \alpha_3 \mathbb{1}(h \neq p) + \alpha_7 \mathbb{1}(\ell = \ell^P) + \boldsymbol{\alpha_\Gamma \Gamma} + \mathbb{E}_{\zeta_{\ell'}}[V'(\Omega, \ell; h)] \Big\}; \quad (3)$$

$$\kappa^\ell c = w_S \mathbb{1}(m=1) + wh - \delta^\ell \cdot \max\Big\{0, h - \tau^S \mathbb{1}(m=1) - \tau^{P,m} \mathbb{1}(\ell = \ell^P)\Big\}.$$

The specification thus flexibly allows women to have different preferences for consumption, leisure, and location based on the presence of young children in the household. When young children are present, the agent must also either dedicate time for caring for their children or absorb child care costs, which depend on their current location, the current location's type, and the woman's marital status. The specification ensures that women never pay for childcare costs if they do not work ($h = 0$), and spouses and grandparents are assumed to contribute fixed time transfers to childcare ($\tau^S$ and $\tau^{P,m}$, respectively) if the woman is either married or living in her parent's location. The grandparents' contribution is allowed to vary based on the marital status of the woman. Furthermore, we allow for unobserved heterogeneity in grandparent helpfulness, such that with probability $P_\tau$ the agent's parents will provide time transfers of zero regardless of marital status[17].

## 3.3 Model Solution

The model is solved via backward induction. In each point of the state space, labor force participation is governed by whether the transient component of the wage offer $\varepsilon$ is sufficiently high. We compute cutoff values of $\varepsilon$ for each element in the state space, after which continuation values can be computed by applying the usual type-1 extreme value formula and using the cutoff values in conjunction with properties of the normal distribution to solve for an agent's expected flow utility in the next period. A more detailed description of the

[17]We have also tried including unobserved heterogeneity in grandparents while estimating the actual transfers provided by unhelpful grandparents and estimated transfers of zero directly.

procedure is as follows:

1. Solve for cutoff values of $\varepsilon$ that govern labor force participation for the terminal age-65 period, where continuation values are zero by construction.

2. Using properties of the normal distribution, solve for *expected* utility $\mathbb{E}_{\varepsilon}[V_{65}(\Omega, \ell)]$ following the optimal hours decision in the age-65 state space.

3. Apply the type-1 extreme value formula to construct the agent's expected utility from choosing their optimal next-period location at age 64:

$$\mathbb{E}_{\zeta_{\ell'}}[V'_{64}(\Omega, \ell; h)] = \bar{\gamma} + \log\left(\sum_{\ell'} \exp\left(\beta \sum_{\Omega'} \mathbb{E}_{\varepsilon}[V_{65}(\Omega', \ell')]\mathbf{Pr}(\Omega'|\Omega, h, \ell') - \Delta(\Omega, \ell')\mathbb{1}\{\ell' \neq \ell\}\right)\right)$$

   where $\bar{\gamma}$ is the Euler-Mascheroni constant. This gives continuation values for all possible combinations of state space and labor supply decisions for the age-64 period.

4. With continuation values in hand, compute cutoff values of $\varepsilon$ for the age-64 period.

5. Repeat steps 2-4 through ages 63 to 22, at which point the model is solved.

Algebraic details for solving cutoff values of $\varepsilon$ and expected utility from hours decisions can be found in in Appendix B.

## 3.4 Discussion

We have presented a tractable model of dynamic labor force participation and migration aimed to capture the geographic constraints imposed by childcare costs and grandparent locations that U.S. women face. The presence of grandparents reduces childcare costs thus reservation wages, allowing women to maintain their participation in the labor force and to continue building work experience that will subsequently raise wages for the remainder of their life. However, these benefits only apply if women are located in the same place as their grandparents, which is a notable constraint given the extent to which migration plays a role in wage growth (Kennan and Walker, 2011). We do not impose that location decisions are one-shot as in García-Morán and Kuehn (2017), however: women may leave their parent's location and then move back when they know that a fertility event is imminent.

The assumption that women only have one child at a time means that our model almost certainly understates childcare costs and the potential effects of childcare subsidy policies on labor force participation, experience, and wages. Moreover, the discretization of geography into nine locations means that we may be suppressing the role that geography plays in wages, which may also have implications for our counterfactual policy predictions. Allowing for a richer geographic structure and potentially an urban/rural distinction while retaining computational tractability in the model may be desirable. Extending the model to account for additional unobserved heterogeneity in wages such that the agents as well as their spouses differ in fixed effects and, possibly, location-specific match effects may also be worthwhile (a la Kennan and Walker (2011)).

# 4 Estimation

## 4.1 Data

We use data on women aged 22-35 in the 2000-onward PSID. All women must be observed at age 22 to be included in our sample. The PSID shifted to a bi-annual schedule starting in 1997 — however, in years following 2000, respondents were asked of their income and hours worked for both the preceding year and the year before. Furthermore, if the respondent had moved across states since their most recent interview, they were asked in which year the move was made. This information, combined with detailed marital and childbirth histories for all respondents, allows us to construct yearly data from the biennial survey with minimal assumptions. If a respondent is observed to be living in a different state since their last interview but does not report the year in which they moved, we assume they moved in the same year as their previous interview.[18]

Importantly, the PSID additionally allows for intergenerational linkages, through which we can track the location of the parents of the respondent. When coding the location of one's parents, we use the state of both parents if both parents are living in the same state (which is the case the overwhelming majority of the time). If the parents are living in different states or if the father's location is missing, we use the location of the mother, and we use the location of the father if the mother's location is missing. Parents are assumed to be living in the agent's home location if the location of both the mother and the father are missing.

The PSID also provides information on the year of birth of the first child of all respondents, as well as the birth years of their four youngest children. We use these years to code fertility events for our sample. If a child is born to a woman in year $t$, it is assumed that the

[18] This happens quite infrequently.

Table 8: Observations by Age

| Age | # Observations |
|---|---|
| 22 | 909 |
| 23 | 909 |
| 24 | 909 |
| 25 | 909 |
| 26 | 813 |
| 27 | 741 |
| 28 | 631 |
| 29 | 554 |
| 30 | 479 |
| 31 | 434 |
| 32 | 362 |
| 33 | 297 |
| 34 | 233 |
| 35 | 174 |

**Notes:** Table presents number of individuals observed at each age in PSID analysis sample. See text for details on sample construction.

woman was aware of the impending birth in year $t-1$ — in other words, $f = 1$ in year $t-1$. We limit our sample to women who are coded as either household heads or the spouses of household heads — thus, information about marital transitions and spousal earnings can be easily obtained from household head information for women labeled as spouses.

We categorize the educational attainment of our sample based on their college status at age 35[19]. Earnings in the data are deflated to real 2012 dollars using the PCE deflator. To constrain the measurement error for wages in the data to reasonable levels, we windsorize hourly wages at the bottom at $7.25 per hour and at the top at the 95th percentile. Observations that report positive hours and zero income are dropped. Observations that reported working 30 hours per week or more are coded as full-time workers, while individuals coded as working less than 30 hours per week are coded as non-participants[20]. Individuals that report working more than 5,820 hours in a year are dropped. Finally, we limit our sample to individuals who are observed continuously in the data for at least 4 years.

[19]For 75% of our sample, college attainment at age 22 was the same as college attainment at age 35. To account for delayed graduation, we exclude college graduates age younger than 25 when evaluating the likelihood function.

[20]Given substantial bunching at 0 or 40 hours per week, alternate thresholds for determining labor force participation have little substantive effect on our results.

These restrictions leave us with a sample of 909 women and 8,354 person-year observations. The median woman in our sample is observed for seven years (i.e. up through age 28), but we have 479 women observed through age 30 and 174 observed through age 35 (see Table 8 for a complete tabulation of ages in our analysis sample). Table 9a presents descriptive demographic and economic statistics broken down by age ranges and college attainment, while Table 9b presents migration statistics in our estimation sample with additional breakdowns by college attainment[21]. Women who have earned a college degree by age 35 have children and marry later, work more, and earn more than their non-college-educated counterparts. College-educated women are unsurprisingly also more migratory, with close to twice the share of college-educated women moving at least once in our data compared to women without a college degree. However, women without a college degree appear more inclined to move back to their parent's location than women with a college degree. The presence of spouses and young children appear to be depress migration rates.

## 4.2 Parameters Estimated Outside the Model

We assume a discount rate of $\beta = 0.95$. Child care costs levels $\delta$ for an hour of care are at the division level following our data from Child Care Aware by averaging across states within a division with population weights. Costs of living $\kappa^\ell$ are taken from the American Chamber of Commerce Research Association's Cost of Living Index.[22] The parameters governing spousal wages are taken a comparable PSID sample to our analysis sample. With the assumption that husbands supply labor exogenously and are of the same age and education level as their wives, the these parameters can be estimated directly from Mincerian wage regressions. Since we assume location wage effects to be equal between men and women, this also allows for the recovery of location wage effects $\eta^\ell$, which are again grouped at the division level[23]. Division populations $N^\ell$ come from Census population estimates.

[21] We do not use sample weights when creating these statistics or when estimating our model. Including longitudinal sample weights available in the PSID does little to change our parameter estimates.

[22] The ACCRA index is a weighted average of costs of food, housing utilities, transportation, health care, and miscellaneous goods and services among different metro areas in the United States. State-level indices have been published from 2016-onward by the ACCRA, and a state-level index constructed by Kennan and Walker (2011) for around 1980 is also available. Unsurprisingly, serial correlation in state-level costs of living is very strong (despite being separated by almost 40 years, the correlation of the two aforementioned sets of values is close to 0.8), so we simply take the midpoint of the two while normalizing the cost of living level of Iowa to be zero before averaging by division with population weights.

[23] See Figure A.1 for representations of state-level childcare costs, wage effects, and living costs.

Table 9: Summary Statistics of PSID Estimation Sample

| Sample | All | | College | | Non-College | |
|---|---|---|---|---|---|---|
| Age | 22/23 | 34/35 | 22/23 | 34/35 | 22/23 | 34/35 |
| LFP Rate | 50.00 | 46.93 | 51.61 | 55.13 | 49.28 | 41.83 |
| | (50.01) | (49.97) | (50.02) | (49.90) | (50.01) | (49.43) |
| Years of Experience | 3.01 | 8.92 | 0.23 | 7.44 | 4.25 | 9.84 |
| | (1.91) | (4.15) | (0.42) | (3.88) | (0.43) | (4.05) |
| Hourly Wage | 13.26 | 17.90 | 16.07 | 20.87 | 11.95 | 15.47 |
| | (5.56) | (7.35) | (5.95) | (7.57) | (4.85) | (6.21) |
| Share Married | 43.40 | 57.00 | 45.00 | 73.08 | 42.69 | 47.01 |
| | (49.58) | (49.57) | (49.79) | (44.50) | (49.48) | (50.01) |
| Young Child Present | 50.99 | 30.47 | 26.79 | 38.46 | 61.76 | 25.50 |
| | (50.00) | (46.08) | (44.32) | (48.81) | (48.62) | (43.67) |
| Observations | 1818 | 407 | 560 | 156 | 1258 | 251 |

(a) Demographic and Economic Statistics

| Sample | All | College | Non-College |
|---|---|---|---|
| Annual Migration Rate | 4.35 | 6.63 | 3.2 |
| | (20.39) | (24.88) | (17.61) |
| *With Children* | 3.38 | 4.83 | 2.83 |
| | (18.08) | (21.45) | (16.58) |
| *If Married* | 3.66 | 5.73 | 2.35 |
| | (18.77) | (23.25) | (15.17) |
| Ever Migrated | 25.33 | 40.32 | 17.81 |
| | (43.49) | (49.06) | (38.26) |
| Share of Moves to $\ell^P$ | 29.01 | 26.06 | 32.08 |
| | (45.45) | (44.03) | (46.82) |
| | | | |
| N | 8,354 | 2,765 | 5,589 |

(b) Migration Statistics

**Notes:** Standard deviations in parentheses. Data from 2001-2017 biennial waves of PSID. Table 9a presents demographic statistics for analysis sample, broken down by educational attainment and age at observation. Table 9b presents migration statistics for the estimation sample, broken down by educational attainment. See text for details on sample restrictions.

Table 10: Parameters Estimated Outside the Model

| Parameter | | Value |
|---|---|---|
| Discount rate | $\beta$ | 0.95 |
| Childcare cost levels | $\delta^\ell$ | Various |
| Location wage effects | $\eta^\ell$ | Various |
| Location living costs | $\kappa^\ell$ | Various |
| Location populations | $N^\ell$ | Various |
| Spouse wage, constant | $\beta_{S,0}$ | 2.234 |
| Spouse wage, education | $\beta_{S,1}$ | 0.571 |
| Spouse wage, experience (linear) | $\beta_{S,2}$ | 0.047 |
| Spouse wage, experience (quadratic) | $\beta_{S,3}$ | -0.0007 |
| Spouse wage, fixed effects | $\mu_S^L, \mu_S^H$ | -0.39, 0.39 |

**Notes:** Table reports values of parameters that are estimated outside the model. Columns 1 and 2 describe the parameters and presents their symbolic representation. Column 3 reports parameter values. See text for details on model and sample construction. See Figure A.1 for representations of state-level childcare costs, wage effects, and living costs.

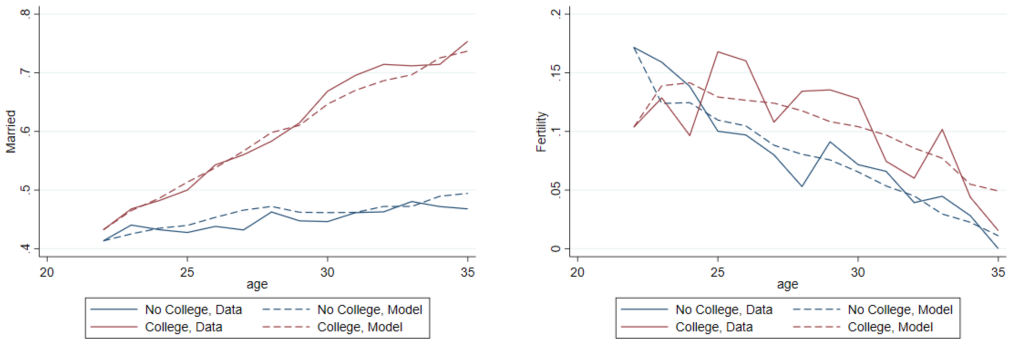
Marriage, divorce, and conception probabilities are estimated via linear probability models that admit as inputs whether the agent is currently married, pregnant, or a parent to young children, as well as a cubic polynomial in age. Probabilities of marital dissolution and formation are also allowed to vary over spousal wage type $\mu_S$. All probabilities are also calculated separately for women with and without a college degree. These linear probability models are estimated directly using our estimation sample. Figure 4 presents the fit of our model with regards to life-cycle profiles of marriage and fertility rates for women with and without a college degree and indicates that our model fits salient features of the data well.

## 4.3 The Likelihood Function

We use maximum likelihood to estimate the remaining parameters of our model. If a woman is never observed to work in the data, we assume that she is a never-working type that always chooses not to work with probability 1. The joint likelihood function for labor force participation, wages, and migration for the $N$ women in our sample, each observed for $T_i$ periods, is given by:

$$L = \prod_{i}^{N} \sum_{\tau} \mathbf{Pr}(\tau) \prod_{t=1}^{T_i} \mathbf{Pr}(h = h_{it}|\Omega_{it}, \ell_{it}) \cdot \mathbf{Pr}(w = w_{it}|\Omega_{it}, \ell_{it}, h_{it}) \cdot \mathbf{Pr}(l' = l'_{it}|\Omega_{it}, \ell_{it}, h_{it}).$$

Figure 4: Model Fit — Marriage and Fertility Life-Cycle Profiles




(a) Marriage (b) Fertility

**Notes:** Figure presents model fit of marriage and fertility rates over lifecycle for women in PSID analysis sample and in data simulated from model. Probabilities estimated separately for women with and without a college degree and depend on marital status, pregnancy, presence of young children, a cubic in age, and spouse wage type. See text for details on sample construction.

We employ a mixture model over unobserved heterogeneity in grandparent transfers, letting $\mathbf{Pr}(\tau)$ denote the probability of the agent being unobserved type $\tau$. The probability of observing wages and hour decisions joint with a location are separable using the assumption that the next-period location shocks are independently distributed from labor supply shocks in each period. For any given element in the state space $(\Omega_{it}, \ell_{it})$, a reservation value of the transient component of the wage offer $\varepsilon^{**}(\Omega_{it}, \ell_{it})$ can be found that governs whether the woman supplies labor in the period[24]. Recall further that wages are measured with error:

$$\log(w) = \beta_0 + \boldsymbol{\beta X} + \eta^{\ell} + \varepsilon + \xi;$$

with $\varepsilon \sim N(0, \sigma_{\varepsilon}^2)$ and $\xi \sim N(0, \sigma_{\xi}^2)$ distributed both i.i.d. and independently from one another. With this assumption, following Eckstein and Wolpin (1989) the first two components of the likelihood function corresponding to labor supply decisions and wages can be defined

[24]For details on deriving these reservation values, refer to Appendix B.

as

$$L = \prod_{i}^{N} \sum_{\tau} \mathbf{Pr}(\tau) \prod_{t=1}^{T_i} \left[ \Phi\left( \frac{\varepsilon^{**}(\Omega_{it}, \ell_{it})}{\sigma_{\varepsilon}} \right) \right]^{1-h_{it}} \cdot$$

$$\left[ \left( 1 - \Phi\left( \frac{\varepsilon^{**}(\Omega_{it}, \ell_{it}) - \rho \frac{\sigma_{\varepsilon}}{\sigma_{\nu}} \nu_{it}}{\sigma_{\varepsilon}\sqrt{1-\rho^2}} \right) \right) \frac{1}{\sigma_{\nu}} \phi\left( \frac{\nu_{it}}{\sigma_{\nu}} \right) \right]^{h_{it}} \cdot$$

$$\mathbf{Pr}(l' = l'_{it} | \Omega_{it}, \ell_{it}, h_{it}),$$

where $\phi$ and $\Phi$ are the standard normal density and cumulative, respectively, $\nu_{it} = \varepsilon_{it} + \xi_{it}$, $\rho = \sigma_{\varepsilon}/\sigma_{\nu}$, and $\sigma_{\nu} = \sqrt{\sigma_{\varepsilon}^2 + \sigma_{\xi}^2}$, leading to $1 - \rho^2$ having the interpretation of the fraction of the wage variance attributable to measurement error. The third component of the likelihood function $\mathbf{Pr}(l' = l'_{it} | \Omega_{it}, \ell_{it}, h_{it})$ can be derived easily following the assumption that the location shocks $\zeta_{\ell'}$ are distributed type-1 extreme value. Denote $V(\Omega, \ell, h, \ell')$ as the expected utility gained from selecting location $\ell'$ following labor supply decision $h$ after starting in state $(\Omega, \ell)$, so:

$$V(\Omega, \ell, h, \ell') = \beta \sum_{\Omega'} \mathbb{E}_{\varepsilon}[V(\Omega', \ell')] \mathbf{Pr}(\Omega' | \Omega, h, \ell') - \Delta \mathbb{1}\{\ell' \neq \ell\}.$$

Recall that $\mathbb{E}_{\varepsilon}[V(\Omega', \ell')]$ represents the expected value of $V(\Omega', \ell')$ after optimizing over the labor supply decision given $\varepsilon$. The method for deriving closed-form expressions of these values is presented in Appendix B, but their recursive nature renders it infeasible to write them out fully. With this, we can now derive the following final representation of the likelihood:

$$L = \prod_{i}^{N} \sum_{\tau} \mathbf{Pr}(\tau) \prod_{t=1}^{T_i} \left[ \Phi\left( \frac{\varepsilon^{**}(\Omega_{it}, \ell_{it})}{\sigma_{\varepsilon}} \right) \right]^{1-h_{it}} \cdot$$

$$\left[ \left( 1 - \Phi\left( \frac{\varepsilon^{**}(\Omega_{it}, \ell_{it}) - \rho \frac{\sigma_{\varepsilon}}{\sigma_{\nu}} \nu_{it}}{\sigma_{\varepsilon}\sqrt{1-\rho^2}} \right) \right) \frac{1}{\sigma_{\nu}} \phi\left( \frac{\nu_{it}}{\sigma_{\nu}} \right) \right]^{h_{it}} \cdot$$

$$\frac{\exp(V(\Omega_{it}, \ell_{it}, h_{it}, \ell'_{it}))}{\sum_{l'} \exp(V(\Omega_{it}, \ell_{it}, h_{it}, \ell')}.$$

## 4.4 Model Assumptions and Identification

The relationship between labor force participation and migration decisions in our model are identified from jointly observing participation, earnings, and location choices for women, conditional on demographic characteristics and location of grand-parents.

First, we assume that the shocks drawn in the model – location preferences, earnings shocks, fertility realization, marriage realization – are all independently and identically distributed across individuals and time. While this may seem a strong assumption at first, we do allow the likelihood of pregnancy and marriage to vary on observable characteristics, including many of the factors that contribute to a woman having a higher or lower earnings potential. This means that this assumption relies only on the weaker assumption that pregnancies and marriages are not correlated across time with the component of earnings that varies idiosyncratically across time. Due to the high rates of unintended and mistimed births in the US (Zolna and Lindberg, 2012), we believe this is a more reasonable assumption. The assumption of independence of location preference shocks and earnings shocks is a stronger assumption: we assume that wage differences across time/individuals are not place-specific and not correlated with amenities in a location in a given year. While the inclusion of amenities along with a reasonably rich wage process in the model helps justify this assumption, allowing for additional heterogeneity in idiosyncratic wage match effects may be helpful as well.

With these distributional assumptions in place, identification of the structural parameters falls out cleanly from the maximum likelihood estimation equation. Following Eckstein and Wolpin (1989) and using equations for reservation wages in Appendix B, the reservation wages $\varepsilon^*$, the wage parameters ($\beta_0, \boldsymbol{\beta}$, and $\mu$), and $\sigma_\varepsilon$ and $\sigma_\xi$ are all identified from data on participation and wages. We can then use the identified $\varepsilon^*$ and our equation for the definition of the reservation wage described in Appendix B to identify $\frac{\alpha_2}{\alpha_1}, \frac{\alpha_e}{\alpha_1}, \frac{\alpha_\mu}{\alpha_1}, \frac{\alpha_c}{\alpha_1}$, and $\frac{\alpha_3}{\alpha_1}$. Based on the similar equation for women with young children, we can identify $\frac{\alpha_6}{\alpha_5}, \frac{\alpha_e}{\alpha_5}, \frac{\alpha_\mu}{\alpha_5}, \frac{\alpha_c}{\alpha_5}, \frac{\alpha_3}{\alpha_5}, \tau_{p1}, \tau_{p0}$, and $\tau_s$. Using any combination of pairs in which the leisure parameter is the same across the presence of children (e.g., $\frac{\alpha_e}{\alpha_1}, \frac{\alpha_\mu}{\alpha_1}, \frac{\alpha_e}{\alpha_5}, \frac{\alpha_\mu}{\alpha_5}$) would allow us to separately identify $\alpha_1$ and $\alpha_5$ and thus separately identify all $\alpha$ parameters governing leisure.

The remaining parameters include the parameters governing preferences for the parent's location, amenities, and the moving cost parameters. These are identified off of the observation of location choices conditional on demographic type and participation in that period. Specifically, we can identify the parent's location preference parameter off the difference in the likelihood of moving to the parent's location $\ell^p$ from some location $k$ and the likelihood of moving to a non-parent's location from that same location $k$ for agents who are similar on all demographic characteristics. The same logic applies for identifying the amenity utility parameters $\boldsymbol{\alpha_\Gamma}$. The moving cost parameters are identified off the differences in likelihood of moving from location $j$ to $k$ versus staying in location $j$ by demographic group. The parameter on population is identified off of the relative likelihood of moving from a small division to a large division vs. from a large division to a small division.

# 5 Results

Parameter estimates are presented in Table 11. When evaluating the likelihood, we exclude women age less than 25 who have a college degree to account for a non-trivial number of women with a college degree who finished their degree in their mid-20s. Standard errors are computed via inverting the numerical Hessian of the likelihood function and taking its diagonal. The estimation recovers preferences for consumption and leisure that increase and decrease respectively over the presence of a small child. The disutility associated from changing one's labor force participation status is substantial, and we also estimate considerable leisure preferences for women with high earnings potential, which is necessary to rationalize their rates of labor force participation that, while higher than low-earning women, are still considerably lower than those of men. The leisure-consumption complementary $\alpha_c$ is positive, reflecting women being less likely to work with higher-earning spouses, all else hold equal. The estimates of wage returns to a college degree and experience are all in line with previous estimates in the literature. We also estimate a meaningful reduction in wages associated with having a child between 0 and 1 year old, but the effect of older children on wages is statistically insignificant.

The estimates of time transfers $\tau$ suggest that spouses and grandparents considerably offset the direct cost of childcare for women with children — indeed, helpful grandparents cover virtually all childcare needs for unmarried women. However, we find substantial heterogeneity in grandparent time transfers based on whether a woman is married, with the married grandparent transfer $\tau^{P,1}$ being quite close 0.4. We note that this is consistent with the heterogeneity in child penalties we estimated in Section 2.2. Moreover, we find that Blacks are more likely to have helpful grandparents than Whites.

Table 11: Parameters Estimated via Maximum Likelihood

| Parameter | | $\hat{\theta}$ | $\hat{\sigma}_\theta$ | $\hat{\theta}$ | $\hat{\sigma}_\theta$ | $\hat{\theta}$ | $\hat{\sigma}_\theta$ |
|---|---|---|---|---|---|---|---|
| *Utility* | | | | | | | |
| Consumption, no children | $\alpha_1$ | 0.103 | 0.011 | 0.097 | 0.014 | 0.112 | 0.029 |
| Leisure, no children | $\alpha_2$ | 1.173 | 0.127 | 1.105 | 0.174 | 1.243 | 0.257 |
| LFP switch penalty | $\alpha_3$ | -0.127 | 0.015 | -0.138 | 0.020 | -0.104 | 0.028 |
| Parent preference, no children | $\alpha_4$ | -0.402 | 0.024 | -0.408 | 0.033 | -0.409 | 0.031 |
| Consumption, with children | $\alpha_5$ | 0.089 | 0.010 | 0.087 | 0.012 | 0.084 | 0.015 |
| Leisure, with children | $\alpha_6$ | 0.765 | 0.086 | 0.742 | 0.127 | 0.724 | 0.050 |
| Parent preference, with children | $\alpha_7$ | -0.405 | 0.104 | -0.046 | 0.204 | -0.634 | 0.120 |
| Consumption/leisure complementarity | $\alpha_c$ | 0.002 | 0.001 | 0.004 | 0.001 | 0.000 | 0.001 |
| College leisure preference modifier | $\alpha_e$ | 0.503 | 0.060 | 0.493 | 0.094 | 0.501 | 0.018 |
| Experience leisure preference modifier | $\alpha_x$ | -0.002 | 0.003 | -0.003 | 0.004 | -0.003 | 0.004 |
| Amenity preference: distance to shore | $\alpha_{\Gamma,1}$ | -0.007 | 0.010 | -0.009 | 0.017 | -0.016 | 0.015 |
| Amenity preference: amenity index | $\alpha_{\Gamma,2}$ | 0.045 | 0.057 | 0.059 | 0.097 | 0.022 | 0.063 |
| Amenity preference: warm days | $\alpha_{\Gamma,3}$ | 0.149 | 0.057 | 0.076 | 0.112 | 0.182 | 0.057 |
| *Time Transfers* | | | | | | | |
| Spouse time transfer | $\tau^S$ | 0.229 | 0.046 | 0.209 | 0.066 | 0.180 | 0.101 |
| Parent time transfer, unmarried | $\tau^{P,0}$ | 0.997 | 0.085 | 0.999 | 0.151 | 0.938 | 0.160 |
| Parent time transfer, married | $\tau^{P,1}$ | 0.388 | 0.077 | 0.394 | 0.105 | 0.460 | 0.204 |
| Probability of $\tau^P = 0$ | $P_\tau$ | 0.687 | 0.032 | 0.708 | 0.054 | 0.658 | 0.056 |
| *Wages* | | | | | | | |
| Wage intercept | $\beta_0$ | 1.972 | 0.020 | 2.000 | 0.031 | 1.928 | 0.007 |
| College effect | $\beta_1$ | 0.458 | 0.016 | 0.440 | 0.022 | 0.488 | 0.024 |
| Experience effect, linear | $\beta_2$ | 0.058 | 0.002 | 0.062 | 0.003 | 0.059 | 0.003 |
| Experience effect, quadratic | $\beta_3$ | -0.002 | 0.000 | -0.002 | 0.000 | -0.002 | 0.000 |
| Child aged 0-1 | $\beta_4$ | -0.085 | 0.016 | -0.082 | 0.021 | -0.069 | 0.028 |
| Child aged 2-4 | $\beta_5$ | -0.028 | 0.015 | -0.026 | 0.021 | -0.028 | 0.024 |
| Wage shock SD | $\sigma_\varepsilon$ | 0.265 | 0.010 | 0.274 | 0.013 | 0.251 | 0.017 |
| Wage measurement error | $\sigma_\xi$ | 0.356 | 0.006 | 0.355 | 0.008 | 0.351 | 0.010 |
| *Moving Costs* | | | | | | | |
| Fixed cost | $\gamma_0$ | 3.840 | 0.208 | 3.804 | 0.265 | 3.906 | 0.326 |
| College effect | $\gamma_1$ | 0.153 | 0.115 | 0.096 | 0.142 | 0.340 | 0.191 |
| Child effect | $\gamma_2$ | 0.277 | 0.138 | 0.261 | 0.173 | 0.422 | 0.250 |
| Marriage effect | $\gamma_3$ | 0.480 | 0.134 | 0.496 | 0.165 | 0.250 | 0.295 |
| Population effect | $\gamma_4$ | 0.007 | 0.054 | -0.022 | 0.068 | 0.042 | 0.092 |
| Sample | | All | | Whites | | Blacks | |
| N | | 8,354 | | 4,837 | | 2,964 | |
| Individuals | | 909 | | 519 | | 324 | |
| Log Likelihood | | -7,429 | | -4,227 | | -2,750 | |

**Notes:** Table presents estimates and standard errors of parameters estimated via maximum likelihood. Data from PSID. See text for details on sample construction and formation of likelihood function.

Of note is that preferences for home location $\alpha_4, \alpha_7$ are estimated to be negative. Since most individuals in our estimation sample start in their parent's location, parameters such as the moving fixed cost $\gamma_1$ are most directly identified from rates of migration out of parent locations. On the other hand, the parent location preference parameters are identified from the rates at which women with and without children *return* to their parent's location. Since parent time transfers are substantial, the model estimates negative values for these preferences to rationalize why we do not see a larger proportion of women moving back to their parents than we would if childcare time transfers were the only factor that influenced the utility from doing so. Amenity preference estimates indicate that the agents in our model prefer higher levels of the Diamond (2016) amenity index, shorter distances to shores, and warmer weather, but only the last of these factors is estimated to be statistically significant.

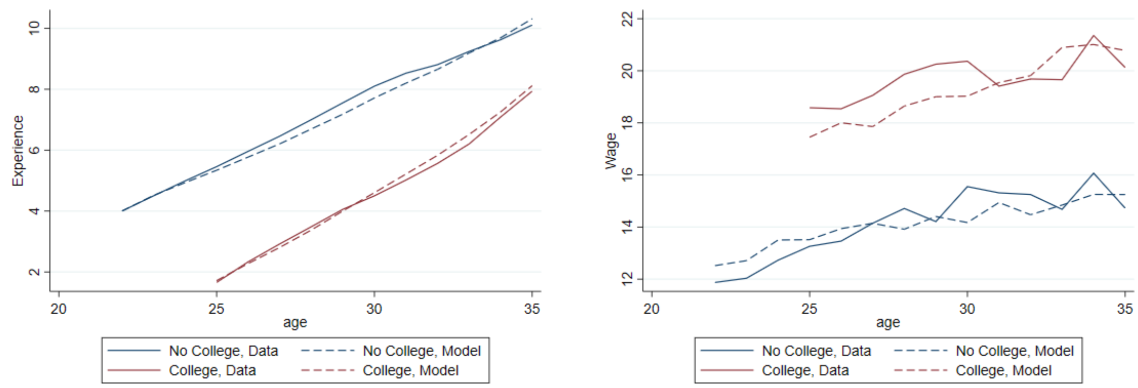
Because utility is linear in consumption, we are able to convert the moving parameters into dollars by dividing by the consumption scaling parameter and then multiplying by the consumption equivalence unit (i.e., one unit of consumption equal \$2000). For the "average" mover, the moving cost is about \$86,000 ignoring the value of the payoffs shocks.[25] For comparison, a woman's life earnings gain would be \$97,000 if, holding all other behavior constant, she moved from the lowest paid region to the highest paid region in age 22 and then stayed in that region for the remainder of her life. Though this is an extreme example of the potential earnings gains from a move, it demonstrates that our moving costs net of payoff shocks are lower than, but of similar scale to the potential earnings gains. However, we will note that these moving costs are the estimated costs for a hypothetical move to an arbitrary location, whereas in the model people will only choose to move to high pay-off locations. Thus, these average costs are higher than the costs that households which actually choose to move will face once pay-off shocks are accounted for.[26]

We estimate our model with our entire analysis sample as well as separately with Blacks and whites, allowing us to compare estimates between the two groups. The primary differences we see between races is in moving costs: we see that Black women have higher moving costs than White women if they have a college degree and if they have a young child. Conversely, they have lower moving costs if married than White women. Our model also shows a

[25] To calculate, we sum $\Delta$ for all individuals who move, discounted by the relevant consumption scaling.That is $\bar{\Delta} = 2000 \times \frac{1}{N_{move}} \sum_{i=1}^{N_{move}} \left[ \left( \frac{\mathbb{1}(a_c \neq \emptyset)_i}{\alpha_5} + \frac{\mathbb{1}(a_c = \emptyset)_i}{\alpha_1} \right) \times (\gamma_0 + \gamma_1 e_i + \gamma_2 \mathbb{1}(a_c \neq \emptyset)_i + \gamma_3 m_i + \gamma_4 N_i^{\ell'}) \right]$.

[26] See Kennan and Walker (2011) for further discussion of the distinction between average moving costs versus average moving costs conditional on moving. Kennan and Walker (2011) show that while the moving costs for households that choose to move to their home location are large, moving costs to non-home locations are actually negative representing the fact that these moves are moves with a large expected future payoffs for the households who make these moves.

Figure 5: Model Fit — Labor Force Lifecycle Profiles




(a) Experience

(b) Hourly Wage

**Notes:** Data from PSID. Figures compare life-cycle trends of experience and wages for women with and without a college degree in estimation sample and data simulated from model. Fit reported for all ages for women with a high school degree and ages 25-onward for women with a college degree. See text for details on sample construction.

larger negative preference for parents if one has a child for Black women than White women. As previously mentioned, this parameter is negative in large part to explain why women do not move home given the high value of being able to use relative care; because Black women are more likely to be single mothers in our data, these time transfers are greater on average for Black women than White women, requiring a larger negative preference parameter to explain why Black women live outside their home location. Other than these parameters, though, the extent of the racial heterogeneity we find is limited.

## 5.1 Goodness of Fit

To assess our model's ability in approximating the true data generating process, we randomly simulate the outcomes of each woman in our estimation sample ten times, starting at age 22 and ending at the final age the given woman is observed in the data, using Bayes' rule to draw unobserved types. We then compare key moments in the estimation sample to those in our simulated data.

Table 12: Model Fit — Labor Force Participation by Location, Marital Status, and Fertility

Panel A: Data

| Marital Status | No Kids | Pregnant | Kids | No Kids, $\ell = \ell^P$ | Pregnant, $\ell = \ell^P$ | Kids, $\ell = \ell^P$ | No Kids, $\ell \neq \ell^P$ | Pregnant,$\ell \neq \ell^P$ | Kids, $\ell \neq \ell^P$ |
|---|---|---|---|---|---|---|---|---|---|
| All | 0.617 | 0.645 | 0.395 | 0.602 | 0.636 | 0.405 | 0.672 | 0.674 | 0.350 |
| $m = 0$ | 0.626 | 0.626 | 0.435 | 0.608 | 0.618 | 0.436 | 0.698 | 0.667 | 0.425 |
| $m = 1$ | 0.568 | 0.558 | 0.386 | 0.556 | 0.543 | 0.398 | 0.624 | 0.609 | 0.332 |

Panel B: Model

| Marital Status | No Kids | Pregnant | Kids | No Kids, $\ell = \ell^P$ | Pregnant, $\ell = \ell^P$ | Kids, $\ell = \ell^P$ | No Kids, $\ell \neq \ell^P$ | Pregnant,$\ell \neq \ell^P$ | Kids, $\ell \neq \ell^P$ |
|---|---|---|---|---|---|---|---|---|---|
| All | 0.633 | 0.515 | 0.380 | 0.628 | 0.518 | 0.397 | 0.654 | 0.503 | 0.293 |
| $m = 0$ | 0.661 | 0.546 | 0.381 | 0.658 | 0.550 | 0.402 | 0.676 | 0.529 | 0.247 |
| $m = 1$ | 0.615 | 0.521 | 0.396 | 0.608 | 0.526 | 0.413 | 0.650 | 0.500 | 0.306 |

**Notes:** Data from PSID. Table compares labor force participation rates for women in estimation sample and data simulated from model. Pregnancy corresponds to woman being pregnant with their first child. See text for details on sample construction.

Table 13: Model Fit — Migration by Fertility

Panel A: Data

| Direction | All | No Kids | Pregnant | Kids |
|---|---|---|---|---|
| $\ell^p$ Out-Migration Rate | 2.06 | 2.69 | 2.69 | 1.52 |
| $\ell^p$ In-Migration Rate | 4.90 | 4.96 | 4.35 | 4.90 |

Panel B: Out of Sample (ACS)

| Direction | All | No Kids | Pregnant | Kids |
|---|---|---|---|---|
| $\ell^p$ Out-Migration Rate | 1.80 | 2.03 | 1.64 | 1.31 |
| $\ell^p$ In-Migration Rate | 4.04 | 4.15 | 4.52 | 3.45 |

Panel C: Model

| Direction | All | No Kids | Pregnant | Kids |
|---|---|---|---|---|
| $\ell^p$ Out-Migration Rate | 1.86 | 2.39 | 1.91 | 1.39 |
| $\ell^p$ In-Migration Rate | 4.47 | 4.70 | 5.19 | 4.14 |

**Notes:** Data from PSID and ACS. Table compares migration rates for women in estimation sample and data simulated from model. Pregnant corresponds to being pregnant with one's first child. See text for details on sample construction.

Figure 5 presents our model's fit in terms of lifecycle profiles of labor market outcomes separately for women with and without a college degree[27]. In general, the model fits the data well, reproducing profiles of wages and experience accumulation that look very similar to the data. The model slightly understates earnings for women with a college degree early in the lifecycle, but by the end of the lifecycle the wage fit is reasonable for both education categories. and slightly overstates rates of migration for women of both levels of education at the beginning and end of the simulation period. The fit of experience suggests that the model does a good job in reproducing the total years worked among women in our sample and overall higher labor force participation rates of college-educated women.

We evaluate the model's fit of labor force participation in more detail in Table 12 by breaking up labor force participation by fertility status (no kids, pregnant with first child, has young children), marital status, and location (in or out of parent's location). Qualitatively, the model can reproduce patterns of women with young children supplying labor markedly less than those without and women with spouses also being less inclined to work than married women. Across all women, the profile of labor force participation the model

[27] We also evaluate the model's fit of similar outcomes for married and unmarried women; see Figure **??**.

outputs over different locations and fertility statuses is reasonable. However, the model does slightly understate the gap in participation between married and single women and understates participation for pregnant women as a whole. Notably, the model also understates participation for single mothers who live outside their parents' location.

Next, we assess the model's fit of migration decisions by breaking down moves according to sending location, destination, and fertility status in Table 13. Since the PSID can have very small samples of movers (particularly pregnant movers), we supplement the table with statistics from the ACS sample used in Section 2 as well. Among all women and women without children, the model predicts sensible rates of migration both out and into the parent location. Moreover, the model is able to qualitatively match the pattern observed in the ACS of pregnant women moving back to their parents' location more frequently, while such behavior is not observed for women with young children in general.

Finally, we evaluate the frequency and quantity of grandparent childcare time transfers observed in our simulated data and compare it to comparable statistics computed from the PSID's Child Development Supplement. In 1997, PSID conducted additional surveys on a sub-sample of 3501 households with children, collecting detailed information on cognitive and non-cognitive skills, health, and care-giving relationships. As part of this survey, they ask parents to report retrospectively about the care-giving arrangements they used throughout the focal child's life, the age at which children first received non-parental care, and whether that care-giver was a non-parent relative or not. Using this data, we tabulate the percent of children who used non-parental relative care exclusively for the child ages 0-4 and compare this to the proportion of families in our model who exclusively use grandparent care rather than mother care or paid child care. In the CDS, 11.5% of families with children 0-4 use relative care compared to 8.4% in our model. While the parents in the CDS are not necessarily the same women in our data[28], the model predicts very sensible rates of relative-based childcare receipt.

# 6 Counterfactual Analysis

Having evaluated our model's performance, we now turn to comparative statics exercises.

[28]In particular, because of the retrospective nature of the 1997 CDS questions, these mothers may be reporting child-care for a child who was 0-4 as early as 1981 which is twenty years prior to the time-frame our model's data covers.

Table 14: Effects of Alternate Demographic Scenarios

Panel A: Impacts of Children only Being Born to Married Women

| Sample | Wages | Wages, $\gamma_1 = \infty$ | Years $x$ | Years $x$, $\gamma_1 = \infty$ | WTP | WTP, $\gamma_1 = \infty$ | # Moves | Share Time with Parents |
|---|---|---|---|---|---|---|---|---|
| All | 0.11 | -0.32 | 0.02 | -0.01 | 23.20 | 15.34 | 0.32 | -0.03 |
| $\tau^P \neq 0$ | -8.18 | -8.16 | -0.50 | -0.51 | 16.99 | 7.28 | 0.37 | -0.04 |
| $\tau^P = 0$ | 3.66 | 3.03 | 0.24 | 0.21 | 25.92 | 18.74 | 0.30 | -0.03 |
| Never SM | -4.96 | -6.42 | -0.24 | -0.30 | 4.37 | 0.00 | 0.17 | -0.02 |
| Ever SM | 4.66 | 5.14 | 0.25 | 0.25 | 40.10 | 29.03 | 0.46 | -0.05 |
| Whites | -1.00 | -2.86 | -0.02 | -0.09 | 12.68 | 4.54 | 0.22 | -0.03 |
| Blacks | 1.53 | 0.79 | 0.01 | -0.03 | 44.29 | 36.43 | 0.43 | -0.04 |

Panel B: Impacts of Removal of Grandparents

| Sample | Wages | Wages, $\gamma_1 = \infty$ | Years $x$ | Years $x$, $\gamma_1 = \infty$ | WTP | WTP, $\gamma_1 = \infty$ | # Moves | Share Time with Parents |
|---|---|---|---|---|---|---|---|---|
| All | -8.37 | -9.88 | -0.49 | -0.56 | -5.34 | -7.18 | 0.03 | -0.01 |
| $\tau^P \neq 0$ | -27.95 | -32.99 | -1.63 | -1.87 | -17.67 | -23.88 | 0.11 | -0.02 |
| $\tau^P = 0$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| Never SM | -5.64 | -7.63 | -0.31 | -0.40 | -2.82 | -4.47 | 0.02 | -0.01 |
| Ever SM | -10.81 | -11.89 | -0.65 | -0.71 | -7.57 | -9.51 | 0.04 | -0.01 |
| Whites | -6.68 | -7.24 | -0.37 | -0.41 | -4.95 | -5.67 | 0.01 | 0.00 |
| Blacks | -11.03 | -12.00 | -0.75 | -0.82 | -6.52 | -8.75 | 0.04 | -0.01 |

**Notes:** SM = Single Mother. The table presents impacts of counterfactual scenarios on mean change in lifetime wages, mean change in years of experience, mean change in number of moves, and mean change in fraction of time spent in parent's location. The second column for wage and experience changes and WTP show results from the counterfactual experiment when moving costs are infinite. Wages in units of $2,000. Results for Whites and Blacks computed using separate parameter estimates shown in Table 11. See text for details on estimation sample and procedure.

## 6.1 Migration and the Family

We begin by predicting lifecycle earnings and migration profiles under alternate demographic scenarios. In a first experiment, we impose that children are only born to married women, that is $P(f = 1|m = 0) = 0$, allowing our model to speak to how recent changes in out-of-wedlock births in the U.S. may have impacted female labor mobility. In a second counterfactual, we evaluate the role of grandparents in wage formation by removing them from the model entirely, setting parent location preferences parameters $\alpha_4, \alpha_7$ as well as grandparent time transfers $\tau^{P,m}$ to zero. Conceptually, the presence of grandparents on wages is ambiguous, since residing with them may increase labor supply and experience in the short run but may also impact wages negatively by discouraging moving to higher-paying locations. In all counterfactuals, we evaluate impacts in the age span of 22 to 55.

Table 14 presents the results of these exercises. For each counterfactual, we estimate impacts in terms of median percent changes in lifetime wages, mean change in years of experience, mean change in number of lifetime moves, and mean change in share of time spent with parents. We also calculate a willingness to pay metric (WTP) by taking the change in utility resulting from the counterfactual scenario and dividing it by $\alpha_1$, the utility scaling parameter for consumption for women without children. We conduct demographic heterogeneity analyses by assessing impacts for women by grandparent helpfulness type and who were or were not ever single mothers in the baseline simulation. We additionally conduct racial heterogeneity analyses for estimating the impacts for whites and Blacks separately while using the separate parameters estimated for them in Table 11. We also compute changes in wages, experience, and WTP in a scenario for moving costs are infinite ($\gamma_1 = \infty$) to assess the importance of accounting for migration when making counterfactual predictions.

The effects of children being born to only married women on wages and experience are generally quite limited. Surprisingly, the effects on work experience are slightly negative, which happens because women acquire experience in part to insure against the state of being a single mother — when the probability of this occurring vanishes, incentives to work decline slightly. While the effects on wages are generally small, we do observe that the counterfactual increases the number of moves made by individuals over the lifecycle by 0.32 on average, with stronger effects for Blacks and women who are ever single mothers, resulting from women who would have had moves encumbered by the presence of children facing smaller moving costs. On a base rate of approximately 1.75 moves over the life cycle in our simulated data, this constitutes an increase of roughly 25%, suggesting that recent increases in single parenthood may have been a contributor to concurrent declines in female labor mobility.

The removal of grandparents is associated with more substantial reductions in wages and experience for women who are ever single parents and Blacks. For example, we see that the existence of grandparents as a potential source of childcare is associated with approximately half a year of experience and subsequently earnings, which corresponds to a percentage increase of about 2.3%. To put these effects in context with the reduced form estimates earlier in the paper, the child penalty gap was about 10 p.p. smaller for mothers living near their grandparents. Removing grandmothers is thus able to account for about 25% of the child penalty we documented, which is reasonable given that not all mothers in our model are living near the grandparents even when they are an available option (and are thus presumably unaffected by the removal of grandmothers) and not all grandparents actually provide childcare assistance.

Additionally, these results demonstrate that parents' mobility is not only influenced by grandparents but also by regional costs for child care: ignoring migration results in overstating the effects of the counterfactual, since when moving is allowed the affected individuals can migrate to areas with lower child care costs or higher wages as a means of insurance. As such, the utility cost of the counterfactual is greater when moving is prohibited, and the difference in utility between the world where moving is allowed is larger for more-affected groups. Eliminating the pull of the parent location, however, does result in increased migration for the same groups of women who see the largest declines in earnings, suggesting that they are substituting from staying in their home location towards either higher paying or lower childcare cost locations when they can no longer take advantage of free relative care in their parent location.

## 6.2 Childcare Subsidies

As a final exercise, we conduct counterfactuals where we halve and then remove child care costs entirely, while breaking down our counterfactual effects by demographics, race, and migration cost scenarios as before. The results of this exercise can be found in Table 15.

Table 15: Effects of Childcare Subsidies

Panel A: Impacts of Halving Childcare Costs

| Sample | Wages | Wages, $\gamma_1 = \infty$ | Years $x$ | Years $x$, $\gamma_1 = \infty$ | WTP | WTP, $\gamma_1 = \infty$ | # Moves | Share Time with Parents |
|---|---|---|---|---|---|---|---|---|
| All | 19.96 | 22.06 | 1.21 | 1.29 | 16.80 | 16.12 | 0.01 | 0.00 |
| $\tau^P \neq 0$ | 7.77 | 9.40 | 0.47 | 0.55 | 8.64 | 6.80 | 0.04 | -0.01 |
| $\tau^P = 0$ | 25.17 | 27.46 | 1.52 | 1.61 | 20.29 | 20.10 | 0.00 | 0.00 |
| Never SM | 14.87 | 17.22 | 0.88 | 0.99 | 12.52 | 11.84 | 0.00 | 0.00 |
| Ever SM | 24.52 | 26.39 | 1.50 | 1.56 | 20.58 | 20.00 | 0.03 | 0.00 |
| Whites | 20.79 | 25.73 | 1.14 | 1.36 | 18.14 | 17.73 | 0.01 | 0.002 |
| Blacks | 18.89 | 19.10 | 1.26 | 1.26 | 14.02 | 12.86 | 0.03 | 0.00 |

Panel B: Impacts of Removing Childcare Costs

| Sample | Wages | Wages, $\gamma_1 = \infty$ | Years $x$ | Years $x$, $\gamma_1 = \infty$ | WTP | WTP, $\gamma_1 = \infty$ | # Moves | Share Time with Parents |
|---|---|---|---|---|---|---|---|---|
| All | 26.82 | 29.93 | 1.72 | 1.86 | 28.35 | 26.60 | 0.03 | -0.01 |
| $\tau^P \neq 0$ | 11.39 | 13.08 | 0.72 | 0.79 | 17.96 | 11.84 | 0.10 | -0.03 |
| $\tau^P = 0$ | 33.41 | 37.13 | 2.15 | 2.31 | 32.72 | 33.01 | 0.00 | 0.00 |
| Never SM | 20.28 | 23.69 | 1.27 | 1.42 | 21.55 | 19.90 | 0.01 | -0.01 |
| Ever SM | 32.67 | 35.51 | 2.12 | 2.24 | 34.47 | 32.62 | 0.05 | -0.01 |
| Whites | 28.49 | 36.50 | 1.64 | 1.97 | 30.72 | 29.90 | 0.02 | -0.01 |
| Blacks | 25.26 | 25.59 | 1.79 | 1.79 | 23.57 | 21.07 | 0.07 | -0.01 |

**Notes:** SM = Single Mother. The table presents impacts of counterfactual scenarios on mean change in lifetime wages, mean change in years of experience, mean change in number of moves, and mean change in fraction of timespent in parent's location. The second column for wage and experience changes and WTP show results from the counterfactual experiment when moving costs are infinite. Wages in units of \$2,000. Results for Whites and Blacks computed using separate parameter estimates shown in Table 11. See text for details on estimation sample and procedure.

In all cases, these policies increase years of experience, labor mobility, and lifetime wages, with particularly strong effects for women who are ever single parents and larger effects for the complete removal of child care costs than halving them. Fully subsidizing childcare increases the lifetime earnings of women by about \$27 (with monetary units in the model scaled by \$2,000, this corresponds to around a \$54,000 impact), which on a basis of \$334 average lifetime earnings in the simulated data corresponds to an approximately 8% increase. For comparison, in the reduced form estimates, we saw that the child penalty for women living in low child care cost states was about 10% lower than for women in high cost states. These wage effects are stronger for women who are single mothers than for women who never are, with percentage wage impacts of around 10% and 5.8% respectively.

Among all women in our sample, the complete removal of childcare costs raises lifetime moves by 0.03, or roughly 2 percent. Similar to before, the effects on earnings are concentrated among women who are at any point single mothers (for whom the percentage effect is closer to 3 percent), and the migration effects are particularly strong for single mothers and Blacks. The effects on earnings and wages are stronger for women who do not have helpful grandparents, since grandparent childcare does crowd out the labor force effects of the policies. However, the migration effects are largest for women who *do* have helpful grandparents, since it is these women for whom the geographic constraint induced by grandparent childcare applies.

A key feature of our results is that they demonstrate that ignoring labor mobility may misstate certain effects of childcare policies. Notably, the effects of the policies on experience and wages are typically larger in the version of the model where moving is prohibited — while this may seem counterintuitive, this happens because women in the no-moving world choose to move to locations that induce labor force participation, such as higher-paying locations or locations with grandparents. This depresses labor force participation in the baseline world with no moving, leaving more room for improvement from the counterfactual policies. Overall, though, the wage effects are fairly comparable regardless of whether moving is allowed.

However, ignoring migration results in the welfare effects of the policies being consistently understated. Across all individuals, the average willingness to pay for the full removal of childcare costs is approximately $\$56,700$, which compared to the average cost in our model of full-time childcare for five years ($\$50,000$) suggests that the policies on a whole may be welfare improving. Unsurprisingly, the willingness to pay for the policies is considerably higher for women that benefit more from them, such as single mothers, Whites (who, recall, have fewer helpful grandparents than Blacks), and women with parents who do not provide childcare. At the same time, the migration mechanism matters more for the groups that more often receive grandparent assistance — compared to the WTP in the world without migration, allowing for migration increases WTP by over 50% for women with helpful grandparents, while it decreases WTP by about 1% for women without helpful grandparents. The same qualitative pattern holds for Blacks and Whites.

While fertility in our model is exogenous, we also assess the sensitivity of these results to allowing a simple fertility elasticity in response to the policies — in particular, we assume that fertility increases by 10% following the halving of childcare costs and by 20% with the full removal, following an elasticity of fertility to reductions in childcare costs estimated by Haan and Wrohlich (2011)[29]. Table A.1 presents the results of this robustness check, reassuringly finding comparable effects of the policies on experience and wages (often larger than in our baseline, since the heightened probability of children in the future appears to induce agents to work more in the present). The additional children result in decreased migration, which results in the WTP for the policies being lower when migration is allowed compared to when not, but this is a mechanical result that should be interpreted with caution, since it would clearly not apply if fertility were endogenous.

# 7 Conclusion

This paper studies how child care costs, the location of extended family, and fertility events influence both the labor force attachment and labor mobility of women in the United States. Using both empirical evidence and a dynamic structural model, we argue that the draw of intergenerational time transfers results in substantial geographic constraints for women with children, thus resulting in children impacting both whether and where women work. As a result, focusing only on labor force participation is likely to understate the potential impacts of childcare policies on wages and welfare of U.S. women. Heterogeneity analyses in our model suggest that accounting for migration is particularly important for women who

[29]Though note that other entries in the literature do not estimate such an elasticity — Bick (2016), for instance, finds that such subsidies do not increase fertility at all. Guner et al. (2020) also model fertility as exogenous, arguing: "We doubt that the inclusion of endogenous parental choices in the analysis could change our quantitative findings in a significant way. Specifically on fertility, child related policies that lead to higher participation rates are unlikely to alter parental decisions. There are countervailing effects that are expected to cancel each other out. Childcare costs are only a small fraction of the lifetime costs of raising children, and a reduction in these costs is balanced by increases in tax rates needed to finance the expansion of childcare subsidies."

are ever single parents and for Blacks. With the COVID-19 crisis introducing substantial upheaval into childcare markets, these geographic constraints may have markedly increased in their importance in recent years.

While the dimensionality reductions in our model result in a high degree of tractability, they do involve considerably suppressing geographic heterogeneity in childcare costs and wage effects. Enriching the geography of the model would be highly desirable in that it would allow us to better capture the extent to which women forego opportunities in stronger labor markets in the face of fertility events. Investigating how these decisions differ among women in and out of urban environments would also be interesting. Further enriching the wage process in the model to account for additional unobserved heterogeneity would also be highly desirable. Another important limitation in our framework is our assumption that fertility is exogenous. Allowing for partial or complete fertility control among women in our framework would allow us to study additional behavioral responses to childcare policies, though robustness exercises suggest that such responses may be small. However, these issues may offer promising avenues for future research.

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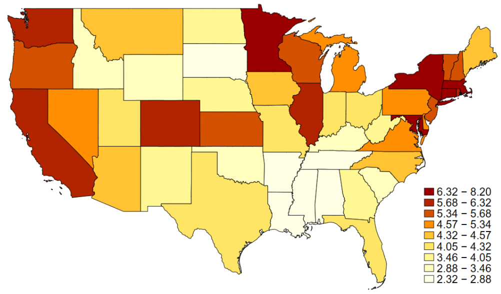
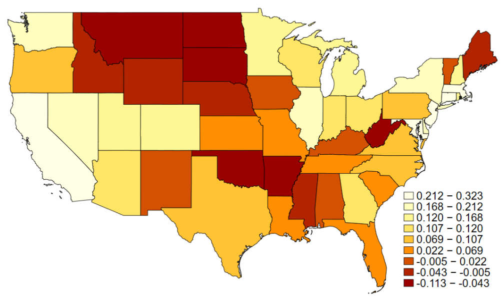
# A Supplementary Figures and Tables

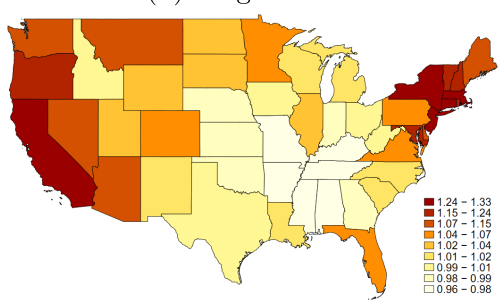
Figure A.1: State Characteristics




(a) Childcare Costs




(b) Wage Effects




(c) Costs of Living

**Notes:** Data on childcare costs from Child Care Aware 2017 report. Units measured in 1000s of 2012 dollars. Wage effects from Mincerian regressions for men in American Community Survey, and costs of living from ACCRA. Wage effect and living cost for Iowa normalized to 0 and 1.

Table A.1: Effects of Childcare Subsidies with Fertilty Response

Panel A: Impacts of Halving Childcare Costs

| Sample | Wages | Wages, $\gamma_1 = \infty$ | Years $x$ | Years $x$, $\gamma_1 = \infty$ | WTP | WTP, $\gamma_1 = \infty$ | # Moves | Share Time with Parents |
|---|---|---|---|---|---|---|---|---|
| All | 28.82 | 32.32 | 1.75 | 1.90 | 9.51 | 15.83 | -0.24 | 0.02 |
| $\tau^P \neq 0$ | 16.04 | 19.01 | 0.95 | 1.09 | 0.49 | 6.31 | -0.20 | 0.02 |
| $\tau^P = 0$ | 34.28 | 38.00 | 2.09 | 2.25 | 13.30 | 19.90 | -0.25 | 0.03 |
| Never SM | 25.50 | 29.55 | 1.52 | 1.71 | 5.15 | 10.87 | -0.25 | 0.02 |
| Ever SM | 31.79 | 34.79 | 1.95 | 2.08 | 13.30 | 20.29 | -0.22 | 0.03 |
| Whites | 31.52 | 45.40 | 1.73 | 2.36 | 25.46 | 78.25 | -0.10 | -0.03 |
| Blacks | 26.22 | 25.83 | 1.67 | 1.72 | 2.95 | 36.25 | -0.22 | 0.04 |

Panel B: Impacts of Removing Childcare Costs

| Sample | Wages | Wages, $\gamma_1 = \infty$ | Years $x$ | Years $x$, $\gamma_1 = \infty$ | WTP | WTP, $\gamma_1 = \infty$ | # Moves | Share Time with Parents |
|---|---|---|---|---|---|---|---|---|
| All | 33.82 | 37.10 | 2.16 | 2.30 | 18.74 | 24.37 | -0.26 | 0.02 |
| $\tau^P \neq 0$ | 19.27 | 21.41 | 1.20 | 1.29 | 7.96 | 9.61 | -0.18 | 0.00 |
| $\tau^P = 0$ | 40.03 | 43.81 | 2.57 | 2.74 | 23.40 | 30.58 | -0.29 | 0.02 |
| Never SM | 29.99 | 33.75 | 1.87 | 2.05 | 11.75 | 16.80 | -0.30 | 0.02 |
| Ever SM | 37.24 | 40.11 | 2.41 | 2.54 | 25.05 | 31.07 | -0.23 | 0.02 |
| Whites | 37.25 | 51.06 | 2.16 | 2.80 | 38.97 | 91.75 | -0.13 | -0.05 |
| Blacks | 30.83 | 30.39 | 2.08 | 2.13 | 8.39 | 41.61 | -0.23 | 0.04 |

**Notes:** SM = Single Mother. Table presents average impacts of counterfactual scenarios on lifetime wages, years of experience, number of moves, and fraction of time spent in parent's location. Second columns for wage and experience changes and WTP conduct counterfactual experiment when moving costs are infinite. Monetary units scaled by \$2,000. Results for whites and Blacks computed using separate parameter estimates shown in Table 11. Fertility probability increases by 10% in halved childcare cost counterfactual and 20% in full subsidy counterfactual. See text for details on estimation sample and procedure.

# B Model Solution Details

This section details the procedure for computing reservation levels of transient wage components $\varepsilon$ and expected value functions when solving the model using the backward induction method described in Section 3.3. There are three stages of life in which households are making decisions: the post-children period (41-65), the post-fertility period (36-40), and the fertility period. The agent's in our model can move at any point and the marital state at age 35 is assumed to be maintained for the remainder of the life-cycle. We focus on reservation wages for age 65 and 64 here to build intuition about decisions without young children and for age 39 for decisions with young children; the procedure for earlier ages is identical after accounting for uncertainty over realizations of marriage and fertility shocks.

In the post-children period, the households no longer ever have young children (i.e., $a_c = \emptyset$). The agent makes a decision of whether the wife should work or not work, which will depend on the realizations of the $\varepsilon$ shock. At age 65, given the other elements of the state space $\Omega$, flow utility from working and not working $u^1_{65}, u^0_{65}$ is given by

$$u^1_{65}(\Omega, \ell) = \alpha_1(w_S(\Omega, \ell)) + \alpha_1 \exp(\beta_0 + \beta_1 e + \beta_2 x + \beta_3 x^2 + \mu + \eta^\ell + \varepsilon) + \alpha_3 \mathbb{1}(p = 0) + \alpha_4 \mathbb{1}(\ell = \ell^P)$$

$$u^0_{65}(\Omega, \ell) = (\alpha_1 + \alpha_c)(w_S(\Omega, \ell)) + \alpha_2 + \alpha_e e + \alpha_x x + \alpha_\mu \mathbb{1}\{\mu = \mu^H\} + \alpha_3 \mathbb{1}(p = 1) + \alpha_4 \mathbb{1}(\ell = \ell^P)$$

$w_S(\Omega, \ell)$ is the realization of spousal income, conditional on spousal characteristics contained within $(\Omega, \ell)$. $p$ is the participation decision the previous period which determines if the person receives the switching cost $\alpha_3$. We ignore costs of living differences and amenities in this formulation, but these can easily be accounted for by dividing $\alpha_1$ and $\alpha_c$ by the relevant $\kappa^\ell$ or by adding the relevant $\boldsymbol{\alpha_\Gamma}$.

Because agents are finitely lived, $V_{66} = 0$ and the value function in the terminal period is then

$$V_{65}(\Omega_{65}, \ell) = \max\left\{u^1_{65}(\Omega, \ell), u^0_{65}(\Omega, \ell)\right\}$$

Participation is governed by the following:

$h = 1$ if

$$\varepsilon_{65} > \log\left(\frac{\alpha_2 + \alpha_e e + \alpha_x x + \alpha_c w_S(\Omega,\ell) + \alpha_3\big(\mathbb{1}(p=1) - \mathbb{1}(p=0)\big)}{\alpha_1}\right) - \underbrace{\left(\beta_0 + \beta_1 e + \beta_2 x + \beta_3 x^2 + \eta^{\ell}\right)}_{G_{65}(\Omega,\ell)} \equiv \varepsilon_{65}^{**}(\Omega,\ell)$$

$h = 0$ otherwise

We can then use this decision rule to calculate the expected utility following the optimal age-65 labor supply choice. That is,

$$\begin{aligned}
\mathbb{E}_{\varepsilon}[V_{65}(\Omega,\ell)] &= \mathbf{Pr}\big(\varepsilon_{65} > \varepsilon_{65}^{**}\big)\mathbb{E}[u_{65}^1(\Omega,\ell)|\varepsilon_{65} > \varepsilon_{65}^{**}] + \mathbf{Pr}\big(\varepsilon_{65} < \varepsilon_{65}^{**}\big)\mathbb{E}[u_{65}^0(\Omega,\ell)|\varepsilon_{65} < \varepsilon_{65}^{**}] \\
&= \alpha_1 w_S + \alpha_4\mathbb{1}(\ell = \ell^P) \\
&\quad + \mathbf{Pr}\big(\varepsilon_{65} > \varepsilon_{65}^{**}\big)\alpha_3\mathbb{1}[p=0] \\
&\quad + \mathbf{Pr}\big(e^{\varepsilon_{65}} > e^{\varepsilon_{65}^{**}}\big)\alpha_1\mathbb{E}(e^{\varepsilon}|e^{\varepsilon_{65}} > e^{\varepsilon_{65}^{**}})\exp(G_{65}(\Omega,\ell)) \\
&\quad + \mathbf{Pr}\big(\varepsilon_{65} < \varepsilon_{65}^{**}\big)\big(\alpha_2 + \alpha_e e + \alpha_x x + \alpha_c w_S(\Omega,\ell) + \alpha_3\mathbb{1}[p=1]\big) \\
&= \alpha_1 w_S + \alpha_4\mathbb{1}(\ell = \ell^P) \\
&\quad + \left[1 - \Phi\left(\frac{\varepsilon_{65}^{**}}{\sigma_{\varepsilon}}\right)\right]\alpha_3\mathbb{1}[p=0] \\
&\quad + \left[1 - \Phi\left(\frac{\varepsilon_{65}^{**} - \sigma_{\varepsilon}^2}{\sigma_{\varepsilon}}\right)\right]\alpha_1 e^{0.5\sigma_{\varepsilon}^2 + G_{65}(\Omega,\ell)} \\
&\quad + \Phi\left(\frac{\varepsilon_{65}^{**}}{\sigma_{\varepsilon}}\right)\big(\alpha_2 + \alpha_e e + \alpha_x x + \alpha_c w_S(\Omega,\ell) + \alpha_3\mathbb{1}[p=1]\big).
\end{aligned}$$

Moving back to period 64, the agent will end the period by realizing their location preference shocks and choosing their optimal age-64 location, $\ell'$, conditional on their current state ($\Omega$), the participation decision made at the beginning of period 64 ($h$), and their current location

($\ell$):

$$\ell' = \arg\max_{k\in\mathbb{N}^{\ell}} \left( \mathbb{1}(k \neq \ell) \times \underbrace{\left(\gamma_0 + \gamma_1 e + \gamma_3 m + \gamma_4 N^k\right)}_{\Delta_k} + \beta \left( \mathbb{E}_{\varepsilon} \sum_{\Omega'} [V_{65}(\Omega', k)] \mathbf{Pr}(\Omega'|\Omega, h, k) \right) + \zeta_k \right)$$

With the assumption that these shocks are drawn from the type-1 extreme value location with a variance normalized to 1, we can calculate the probability of choosing location $\ell'$:

$$Pr(\ell_{64} = \ell'|\Omega, \ell, h) = \frac{\exp\left(\mathbb{1}(\ell' \neq \ell) \times \Delta_{\ell'} + \beta \left(\mathbb{E}_{\varepsilon} \sum_{\Omega'} [V_{65}(\Omega', \ell')] \mathbf{Pr}(\Omega'|\Omega, h, \ell')\right)\right)}{\sum_k \exp\left(\mathbb{1}(k \neq \ell) \times \Delta_k + \beta \left(\mathbb{E}_{\varepsilon} \sum_{\Omega'} [V_{65}(\Omega', k)] \mathbf{Pr}(\Omega'|\Omega, h, k)\right)\right)}$$

and the expected utility following the optimal decision as:

$$\mathbb{E}_{\zeta_{\ell'}}[V'_{64}(\Omega, \ell; h)] = \bar{\gamma} + \log\left( \sum_{\ell'} \exp\left( \beta \sum_{\Omega'} \mathbb{E}_{\varepsilon}[V_{65}(\Omega', \ell')] \mathbf{Pr}(\Omega'|\Omega, h, \ell') - \Delta_{\ell'} \mathbb{1}\{\ell' \neq \ell\} \right) \right).$$

This then allows us to express the age-64 value function as:

$$V_{64}(\Omega, \ell) = \max\left\{ u^1_{64}(\Omega, \ell) + \mathbb{E}_{\zeta_{\ell'}}[V'_{64}(\Omega, \ell; 1)], u^0_{64}(\Omega, \ell) + \mathbb{E}_{\zeta_{\ell'}}[V'_{64}(\Omega, \ell; 0)] \right\},$$

where $u^1_{64}(\Omega, \ell)$ and $u^0_{64}(\Omega, \ell)$ are defined comparably to their age-65 counterparts. The decision rule for working given $\varepsilon_{64}$ is then given by:

$h = 1$ if

$$\begin{aligned}\varepsilon_{64} > \log\Bigg( &\frac{\alpha_2}{\alpha_1} + \frac{\alpha_e}{\alpha_1} e + \frac{\alpha_x}{\alpha_1} x + \frac{\alpha_c}{\alpha_1} w_S(\Omega, \ell) + \frac{\alpha_3}{\alpha_1}\left(\mathbb{1}(p=1) - \mathbb{1}(p=0)\right) \\ &+ \frac{1}{\alpha_1} \mathbb{E}_{\zeta_{\ell'}}\left([V'_{64}(\Omega, \ell; 0)] - \mathbb{E}_{\zeta_{\ell'}}[V'_{64}(\Omega, \ell; 1)]\right)\Bigg) - G_{64}(\Omega, \ell) \equiv \varepsilon^{**}_{64}(\Omega, \ell)\end{aligned}$$

$h = 0$ otherwise.

This then allows us to express the expected utility following the age-64 labor supply choice as follows:

$$
\begin{aligned}
\mathbb{E}_{\varepsilon}[V_{64}(\Omega,\ell)] \quad = \quad & \alpha_1 w_S + \alpha_4 \mathbb{1}(\ell = \ell^P) \\
& + \left[1 - \Phi\left(\frac{\varepsilon_{64}^{**}}{\sigma_\varepsilon}\right)\right]\Big(\alpha_3 \mathbb{1}[p=0] + \mathbb{E}_{\zeta_{\ell'}}[V'_{64}(\Omega,\ell;1)]\Big) \\
& + \left[1 - \Phi\left(\frac{\varepsilon_{64}^{**} - \sigma_\varepsilon^2}{\sigma_\varepsilon}\right)\right]\alpha_1 e^{0.5\sigma_\varepsilon^2 + G_{64}(\Omega,\ell)} \\
& + \Phi\left(\frac{\varepsilon_{64}^{**}}{\sigma_\varepsilon}\right)\Big(\alpha_2 + \alpha_e e + \alpha_x x + \alpha_c w_S(\Omega,\ell) + \alpha_3 \mathbb{1}[p=1] + \mathbb{E}_{\zeta_{\ell'}}[V'_{64}(\Omega,\ell;0)]\Big),
\end{aligned}
$$

which in turn allows us to compute age-63 continuation values, and so on. This continues recursively in the same fashion until we reach age 39, which is the last year in which an agent may have a young child. For those without children at 39, the decision process is unchanged. For those with a child, they now have the costs of child care to consider in their hours decision and the location of parents as a source of cheaper care to consider in their location decision.

For a person with a young child, given the other elements of the state space $\Omega$, flow utility from working and not working $u^1, u^0$ is given by:

$$
\begin{aligned}
u^1(\Omega,\ell,a_c \neq \emptyset) \quad = \quad & \alpha_5\Big(w_S(\Omega,\ell) + \exp\big(\beta_0 + \beta_1 e + \beta_2 x + \beta_3 x^2 + \eta^\ell + \varepsilon\big) \\
& - \delta_\ell\big(1 - \tau_{pm}\mathbb{1}(\ell = \ell^P) - \tau_s m\big)\Big) + \alpha_3 \mathbb{1}(p=0) + \alpha_7 \mathbb{1}(\ell = \ell^P); \\
u^0(\Omega,\ell,a_c \neq \emptyset) \quad = \quad & (\alpha_5 + \alpha_c)(w_S(\Omega,\ell)) + \alpha_6 + \alpha_e e + \alpha_x x + \alpha_3 \mathbb{1}(p=1) + \alpha_7 \mathbb{1}(\ell = \ell^P).
\end{aligned}
$$

At age 39, the expected utility following the optimal location decision, $\mathbb{E}_{\zeta_{\ell'}}[V'_{39}(\Omega,\ell;h)]$ follows the same general form as the previously described expected utility in period 64. This, combined with the flow utility, gives us the following participation decision rule:

$h = 1$ if

$$
\begin{aligned}
\varepsilon_{39} > \log\Bigg(& \frac{\alpha_6}{\alpha_5} + \frac{\alpha_e}{\alpha_5}e + \frac{\alpha_x}{\alpha_5}x + \frac{\alpha_c}{\alpha_5}w_S(\Omega,\ell) + \frac{\alpha_3}{\alpha_5}\big(\mathbb{1}(p=1) - \mathbb{1}(p=0)\big) \\
& + \frac{1}{\alpha_5}\mathbb{E}_{\zeta_{\ell'}}\big([V'_{39}(\Omega,\ell;0)] - \mathbb{E}_{\zeta_{\ell'}}[V'_{39}(\Omega,\ell;1)]\big) + \delta_\ell\big(1 - \tau_{pm}\mathbb{1}(\ell = \ell^P) - \tau_s m\big)\Big)\Bigg) - G_{39}(\Omega,\ell) \equiv \varepsilon_{39}^{**}(\Omega,\ell);
\end{aligned}
$$

$h = 0$ otherwise.

There are two notable differences in the reservation wage for women with children relative to those without. First, the parameters governing valuation of consumption ($\alpha_5$) and the value of leisure ($\alpha_6$) differ, potentially raising the reservation wage relative to non-mothers if $\alpha_6$ is higher than $\alpha_2$ or lowering the reservation wage if $\alpha_5$ is higher than $\alpha_1$. Second, the added cost of child care $\delta^\ell$ will raise the reservation wage for mothers.

This then allows us to express the expected utility following the age-64 labor supply choice as follows:

$$\begin{aligned}
\mathbb{E}_\varepsilon[V_{39}(\Omega,\ell)] \quad = \quad & \alpha_5 w_S + \alpha_7 \mathbb{1}(\ell = \ell^P) \\
& + \left[1 - \Phi\left(\frac{\varepsilon_{39}^{**}}{\sigma_\varepsilon}\right)\right]\Big(\alpha_3 \mathbb{1}[p=0] - \alpha_5 \delta_\ell \big(1 - \tau_{pm}\mathbb{1}(\ell = \ell^P) - \tau_s m\big) + \mathbb{E}_{\zeta_{\ell'}}[V'_{39}(\Omega,\ell;1)]\Big) \\
& + \left[1 - \Phi\left(\frac{\varepsilon_{29}^{**} - \sigma_\varepsilon^2}{\sigma_\varepsilon}\right)\right]\alpha_5 e^{0.5\sigma_\varepsilon^2 + G_{39}(\Omega,\ell)} \\
& + \Phi\left(\frac{\varepsilon_{39}^{**}}{\sigma_\varepsilon}\right)\Big(\alpha_6 + \alpha_e e + \alpha_x x + \alpha_c w_S(\Omega,\ell) + \alpha_3 \mathbb{1}[p=1] + \mathbb{E}_{\zeta_{\ell'}}[V'_{39}(\Omega,\ell;0)]\Big),
\end{aligned}$$

which in turn allows us to compute age-38 continuation values. The location choice decision at age 38 takes the same form as previously, though with the addition of a component of the moving cost for parents with a young child and with an additional component of expected future utility (i.e., the future child care costs) varying by location. The process continues in this manner until age 35, at which fertility shocks and marriage shocks enter the model. As these shocks merely impact the probability of being in a given state $\Omega$ in the following period, we omit discussion of life-period one decision solutions; the remainder of the model can thus be solved in a similar fashion until reaching age 22.

# C Divisional Groupings of States

- **New England (NE):** Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, Vermont.
- **Mid-Atlantic (MA):** New Jersey, New York, Pennsylvania.
- **East North Central (ENC):** Illinois, Indiana, Michigan, Ohio, Wisconsin.
- **West North Central (WNC):** Iowa, Kansas, Minnesota, Missouri, Nebraska, North Dakota, South Dakota.
- **South Atlantic (SA):** Delaware, Florida, Georgia, Maryland, North Carolina, South Carolina, Virginia, District of Columbia, West Virginia.
- **East South Central (ESC):** Alabama, Kentucky, Mississippi, Tennessee.
- **West South Central (WSC):** Arkansas, Louisiana, Oklahoma, Texas.
- **Mountain (MO):** Arizona, Colorado, Idaho, Montana, Nevada, New Mexico, Utah, Wyoming.
- **Pacific (PA):** Alaska, California, Hawaii, Oregon, Washington.