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Research Note: New Evidence on the Motherhood Wage Penalty

Abstract

U.S. women's age of first birth has increased substantially. Yet, little research has considered how this changing behavior may have affected the motherhood pay penalty, or the wage decrease with a child's arrival, experienced by the current generation. Using Rounds 1-19 of the National Longitudinal Survey of Youth 1997 (NLSY97), we examine shifts in hourly pay with childbirth for a cohort of women who became mothers mostly in the 2000s and 2010s. Results from fixed-effects models indicate that the motherhood pay penalty for NLSY97 women who had their first child before their late 20s is generally similar to that of previous cohorts. Those becoming mothers near or after age 30, however, encounter a parenthood premium like men do. The growing proportion of women delaying motherhood, coupled with the rising heterogeneity in motherhood wage outcomes by childbearing timing, contributes to a comparatively small motherhood penalty for this recent cohort. The pay advantage of late mothers cannot be explained by factors such as their labor market locations, number of children, stage of childrearing, marital status, or ethnoracial composition. Instead, the hourly gain stems from such mothers' tendency to reduce working hours more than other mothers without experiencing a commensurate decrease in total pay. Unlike the fatherhood premium, the premium for late mothers does not lead to a real boost in income.

Keywords: Motherhood penalty, fertility timing, changes over time, working hours, parenthood premium

Family demographers have extensively documented the tendency for women to experience a wage decrease with each child's arrival (Budig and England 2001; Deming 2022; Gough and Noonan 2013). Commonly referred to as the motherhood pay penalty, this decrease contributes to a widening pay gap between women and men throughout their life course, ultimately perpetuating gender inequality (Angelov et al. 2016; England 2005). Research on U.S. mothers' pay penalty to date predominantly relies on data from the National Longitudinal Survey of Youth 1979 (NLSY79), which collects information from a cohort born in 1957-1964 (Amuedo-Dorantes and Kimmel 2005; Budig and England 2001; Budig and Hodges 2010; Doren 2019; England et al. 2016; Gangl and Ziefle 2009; Glauber 2007; Gough 2017; Yu and Hara 2021). A few studies use longitudinal data of women born even earlier (Anderson et al. 2003; Avellar and Smock 2003; Kahn et al. 2014; Taniguchi 1999). Consequently, our knowledge about the magnitude and prevalence of motherhood penalty in the United States is primarily based on women having childbirths in the 1980s-1990s or earlier.

Motivated by the substantive shifts in women's socioeconomic opportunities in recent years, a handful of studies have used data from the Current Population Study (CPS) to examine changes in the motherhood wage penalty through the 2010s (Glauber 2018; Pal and Waldfogel 2016; Weeden et al. 2016). According to these studies, the penalty has declined in recent decades, especially for married women, high earners, and non-Hispanic Whites. Because the CPS generally does not track wage trajectories before and after childbearing, however, the motherhood penalties estimated with such data must rely on comparisons between women who select to be mothers and women who do not. As a result, the observed decline in mothers' penalty could simply reflect changes in the selectivity into motherhood. For example, if women with unobserved characteristics that lead to a higher income (e.g., high productivity) used to be

more likely to forego motherhood but are no longer so, we could find mothers and childless women to converge in earnings over time.

Precisely due to the selectivity issue, most research on the motherhood penalty uses longitudinal data, estimating the penalty by comparing women's earnings before and after childbirth (Budig and England 2001; Budig and Hodges 2010; England et al. 2016; Glauber 2007; Yu and Hara 2021). To our knowledge, only one study applies this approach to a contemporary cohort of women.[1] Specifically, using data from the National Longitudinal Survey of Youth 1997 (NLSY97), Yu and Kuo (2017) show that the magnitude of motherhood wage penalty for women born in 1980-1984, who likely became mothers in the 2000s and 2010s, is comparable to that of their mothers' generation (i.e., NLSY79 women). This finding echoes Avellar and Smock's (2003) that the level of pay decrease with childbearing among NLSY79 women is similar to that observed in an earlier cohort, suggesting the motherhood penalty is highly resistant to social change.

At the time of Yu and Kuo's (2017) study, NLSY97 women were largely in their early 30s. More years of NLSY97 data have become available since then; examining the motherhood penalty with the additional data can potentially alter our understanding. This is because Millennials—the group to which NLSY97 women belong—have been postponing parenthood more than prior generations (Barroso et al. 2020). From 1970 to 2017, women's average age of first birth has increased from 21 to 27 years (Guzzo and Payne 2018). For cohorts undergoing childbearing at an earlier age, such as NLSY79 women, the motherhood pay penalty estimated in women's early 30s hardly differs in magnitude from that at the end of their reproductive years

[1] Jee and colleagues (2019) use longitudinal data from the Panel Survey of Income Dynamics and fixed-effects models to compare the motherhood pay penalty from 1986-2014, but their estimated penalty for the period 2006-2014 is for all women observed during that period, most of whom were born years before the NLSY97 cohort. It is notable that their analysis suggests a fairly stable motherhood penalty over time, just like Yu and Kuo's (2017).

(Budig and England 2001; Yu and Hara 2021). For Millennials, however, missing the experiences of "late mothers," who have their first or second child in their mid- to late 30s, could bias the estimate considerably. Prior research shows that women delaying childbirth face a smaller motherhood pay penalty (Amuedo-Dorantes and Kimmel 2005; Doren 2019; Gough 2017; Taniguchi 1999). With women becoming increasingly heterogeneous in earnings in recent years (Blau and Kahn 2007), the pay penalties for early and late mothers could also have diverged further, making undercounting the latter's childbirths especially problematic. Of course, because mothers' pay disadvantage may decline as their children grow up (Anderson et al. 2003), having additional years of observations may also avoid overestimating this disadvantage with just data from early childrearing years.

In this research report, we use 19 rounds of NLSY97 data (last collected in 2019-2020) to examine the motherhood wage penalty for early Millennials. Our study has three objectives. First, we track this contemporary cohort into its late 30s to more accurately estimate the average penalty for recent mothers. Doing so enables us to adjudicate the contradictory findings between prior cross-sectional and longitudinal studies regarding the stability of mothers' pay disadvantage over time (Glauber 2018; Jee et al. 2019; Pal and Waldfogel 2016; Yu and Kuo 2017). Second, because any observed change in the estimated motherhood penalty with additional NLSY97 rounds is likely attributable to the incorporation of late mothers' experiences, we also investigate how the penalty varies by motherhood timing. Uncovering the extent of this variation helps inform the implications of recent shifts in fertility behavior for the gender gap in pay. Finally, we explore explanations for any potential differences in motherhood penalties by women's age of first birth. Although Studies of earlier cohorts have found little pay penalty for women delaying childbirth until their late 20s (Amuedo-Dorantes and Kimmel 2005;

Doren 2019; Taniguchi 1999), a systematic examination of the reasons behind this tendency is absent. Yet, understanding why late mothers differ is increasingly important, as their experiences will likely contribute to the overall wage cost of childbearing even more in the years to come.

Research on the variations in mothers' pay disadvantage suggests several reasons why the extent of this disadvantage may depend on childbirth timing. First, having had more undivided time for career development, women who become mothers later in life may occupy different labor market positions from other mothers. Research shows that sorting by firm type, occupation, and industry explains much of the gender gap in pay and mothers' pay disadvantage compared to nonmothers (Petersen and Morgan 1995; Yu and Hara 2021). The sorting in the labor market could similarly explain why late mothers may fare better. Second, women who delay motherhood likely have fewer and younger children by their late 30s. Research suggests that mothers' pay disadvantage may vary by parity and the child's age (Abendroth et al. 2014; Anderson et al. 2003; Wilde et al. 2010). The number or age of children at the time of observation could therefore explain the penalty gap between earlier and later mothers. Third, later mothers may be more likely to be married than earlier mothers, which could account for the potential penalty heterogeneity. Nonetheless, prior findings on the variation by marital status are mixed, with some showing a greater penalty and others no difference or a smaller penalty for married mothers (Budig and England 2001; Glauber 2007; Gough and Noonan 2013; Pal and Waldfogel 2016). Thus, it is unclear whether a greater likelihood of being married would necessarily lead to a smaller motherhood penalty.

Fourth, evidence from earlier cohorts indicates a larger motherhood penalty for White than Black or Hispanic women (Glauber 2007; Gough and Noonan 2013). Some suggest that ethnoracial backgrounds and subcultures affect women's anticipated relationship stability and

family roles after childbirth, which in turn shape how they adjust work behavior and how their earnings change with motherhood (Deming 2022). The stereotype about mothers' low work commitment, as demonstrated by Correll and others (2007), could also hurt Black and Hispanic women less because employers see them as less productive than Whites even without children. At the same time, a cross-sectional study indicates the Black-White gap in the motherhood pay penalty may have been reversed in recent decades (Pal and Waldfogel 2016). It is possible that Black mothers have experienced especially drastic increases in relationship instability (Brown et al. 2016), which amplifies the difficulty to simultaneously meet work and childrearing demands. Because White women tend to enter motherhood later than Black or Hispanic women (Mathews and Hamilton 2016), it is important to test whether the ethnoracial differences between earlier and later mothers explain their relative penalties.

Finally, women who delay childbearing may have accumulated more resources to shield themselves from the extra burden brought by motherhood, thus needing to alter work effort less. A smaller change in effort could lead to a lower pay penalty (Gough and Noonan 2013). To gauge work effort, we analyze how women entering motherhood at different times adjust their working hours. Because employers may tie work effort and compensations to the job's normative time demand rather than workers' actual hours (Yu and Kuo 2022), we also investigate whether late mothers experience less change in the normative time demand, and how the modifications of work hours or normative time demand affect their total post-childbirth earnings.

To summarize, using longitudinal data of women who just underwent the typical reproductive years, our analysis addresses how incorporating late mothers' experiences may revise the current estimate of the motherhood penalty; how heterogeneous mothers' pay disadvantage has become; and why the pay penalty varies by childbirth timing for contemporary

mothers. By doing so, we provide up-to-date knowledge of the prevalence and variation of mother's pay disadvantage and shed light on the implications of changing fertility behavior.

Data and Methods

The data for the analysis come from Rounds 1-19 of the NLSY97, which followed a sample of individuals born in 1980-1984 from 1997-2011 annually and biannually afterwards. By Round 19, the women in the sample were mostly in their late 30s, toward the end of their reproductive years.[2] We pooled the rounds together to create person-year data. We selected all person-years of women during which they were at least 16 years old, reported working as an employee, and received a wage. The self-employed was excluded because their earnings may depend on other factors (e.g., capital investment), although including them in an earlier analysis did not affect our main results. We also omitted observations with invalid information about parenthood status (<1%) or other key variables (~5%). The resulting sample contains 46,997 person-years from 4,175 women.

The main outcome for the analysis is hourly pay for the job reported at the interview. We took the natural log of hourly pay to reduce skewness. This transformation also enables us to interpret coefficients as percentage changes in pay. Table 1 presents details for how we measured this and other variables, along with their descriptive statistics. To show how motherhood changes women's time spent at their jobs and their total income, we also fitted models with respondents' log weekly working hours or log weekly pay as the dependent variable.

[Table 1 about here]

[2] The NLSY97 provides a binary indicator of respondents' gender but does not explain how "women" are defined. Because the gender information was collected only in 1997, it is presumably based on respondents' self-identification as male or female during adolescence. For simplicity, we use "women" to refer to those identified as women in the survey hereafter.

Our main predictor is respondents' number of children. We mostly treated this variable as continuous but adopted an alternative specification with categorical dummies (1, 2, or 3 and more children) for some models. To examine how the motherhood penalty differs by childbirth timing, we included the age range of first birth. We followed Taniguchi (1999) to consider women entering motherhood at age 28 or older as late mothers, while those becoming mothers by age 22 and at ages 23-27 as early and average mothers, respectively.[3]

Our models included several measures of human capital: education, school enrollment, work experience, tenure with the current job and tenure squared (to capture the plateauing effect after staying at a job for too long), and number of major employment breaks. We also controlled for region of residence and living in urban areas or not, as wage levels vary by geographic location. To test whether labor market locations explain the potential penalty differences by motherhood timing, we introduced firm size, whether the firm has multiple locations, and the firm's unionization status. We also followed Yu and Hara (2021) and used occupation and industry fixed effects, which can be understood as including one dummy variable for each occupation or industry, to account for all time-constant occupational and industrial characteristics that may affect wages. To examine whether the motherhood penalty varies by women's occupational skill or industry, we further transformed the 3-digit occupational codes into a continuous measure of occupational skill level and created 18 broad industrial categories. To assess the possibility that the different childrearing stages or marital statuses for women with various first-birth timings explain their relative pay penalties, we constructed indicators of the

[3] Using age 28 as the cut-off point means that some late mothers would have been mothers already in Yu and Kuo's (2017) sample, but the extra data rounds may still capture their second or third childbirth. Because the per-child penalty may differ between later and earlier mothers, incorporating more births of the former could alter the average estimate of the motherhood penalty. The results were similar when we tried other age cut-off points for the three groups of mothers.

youngest child's age (if respondents had a child) and marital status. To show whether earlier and later mothers' ethnoracial compositions may be relevant to any penalty differences, we divided respondents by their self-identified race-ethnicity and fitted separate models for each group (characteristics of each group in Table A2 in Online Supplement).

Like prior research on mothers' pay disadvantage (Budig and England 2001; Yu and Kuo 2017), we used two-way fixed-effects models, which can be expressed as:

$$\ln(pay¿¿it) = \gamma_0 + \gamma_1 children_{it} + a_j X_{jit} + \mu_i + S_t + \varepsilon_{it}, ¿$$

where the outcome is the log hourly pay of person i at survey year t; $\gamma_0$is the intercept; γ_1 is the coefficient for the number of children; X_{jit}denotes j time-varying variables that may also affect earnings (e.g., work experience, location); $\mu_i$and S_t are fixed effects for i individuals and t survey rounds in the data, respectively; and $\varepsilon_{it}$is the error term. Because μ_i captures all time-constant unobserved individual characteristics, including personality traits conducive to the selection into motherhood, this model provides a more rigorous estimate of the wage gain or loss linked to motherhood than cross-sectional models. S_k accounts for all year-to-year influences on wages, including inflation and other macroeconomic shifts.

To assess how the motherhood penalty varies by first-birth timing, we further estimated:

$$\ln(pay¿¿it) = \gamma_0 + \gamma_1 children_{it} + \gamma_2 children_{it} \times age_{¿i} + a_j X_{jit} + \mu_i + S_t + \varepsilon_{it}, ¿$$

where $age_{¿i}$ is the age range within which person i had her first birth. This variable's main effect is omitted because μ_i absorbs the effects of all factors that do not vary for individuals over time. To test whether occupational and industrial locations explain the differences in motherhood penalty by childbearing timing (γ_2), we added occupation and industry fixed effects to the right-hand side of the equation in further models. We also included the interactions between $children_{it}$ and occupational skill level, broad industry, and firm characteristics in alternative models. We

similarly estimated models with the interaction between $children_{it}$ and marital status, as well as fitted separate models by race-ethnicity, to observe how γ_2 changes. To investigate how the youngest child's age may explain γ_2, we included the youngest child's age and its interaction with age_{it} in some models. We further fitted a model where $children_{it}$ is measured as a categorical variable (1, 2, or 3 and more children) to evaluate whether a disproportionately large penalty for higher-order children explains the potentially lower penalty for late mothers, who may have just one or two children.

Finally, we estimated fixed-effects models with log weekly working hours and normative time demand as the outcomes to assess how changes in work effort or perceived work effort correspond to earlier and later mothers' penalties. The measure of normative time demand was derived from an indicator of typical working hours from the Occupational Information Network (O*NET), as detailed in Table 1. Because shifts in working hours can affect not just hourly pay but also total wages, we fitted another model predicting weekly pay to grasp the full economic impact of motherhood. All models were estimated with the NLSY97 longitudinal weights and robust standard errors.

Results

Table 2 presents fixed-effects models predicting women's log hourly wages. Model 1 uses the same data as in Yu and Kuo (2017)—through Round 16—and shows that net of human capital and marital status, having a child is associated with a 4.3% decrease in hourly pay, similar to the estimated penalty for women born a generation earlier (Budig and England 2001).[4]

[4] Our estimate was identical to Yu and Kuo's (2017), 4.2%, when we used the initial sampling weights as did that study. We also replicated Budig and England's (2001) model with 16 rounds of NLSY97 data (which cover a similar age range as their study) and yielded an estimate of 4.4%, which is very close to theirs, 4.7% (see Table A3 in Online Supplement).

When including all 19 rounds of data in Model 2, however, the pay penalty reduces by over 40%, with each child tied to just a 2.5% decrease.[5] Model 3 adds the interaction between number of children and age of first birth, making the main effect of number of children represent the wage change for women entering motherhood by age 22. Based on this model, the per-child penalty is 3.4-4% for women having first birth by age 27, but those becoming mothers at age 28 or older experience a 4.1% pay *premium* (-0.034+0.075=0.041; $p < .01$). The difference between late and earlier mothers (>7%) is much larger than the 2-3% documented for a prior generation (Taniguchi 1999). As late mothers' pay premium offsets the penalty faced by other mothers, incorporating more of their childbirths in the data leads to a decrease in the average penalty.[6]

[Table 2 about here]

To shed light on the differences between late and other mothers, Table 3 presents women's characteristics by childbirth timing. Late mothers have fewer children, but a much larger proportion of them (63%) added at least one child since Round 16. These mothers also have more education, longer work experience and job tenure, more extended work hours, higher occupational status, and are more likely to be married and White than all other groups of women, including nonmothers. Consistent with their having the most human capital, late mothers receive the highest pay, followed by nonmothers, average mothers, and early mothers, in that order.

[Table 3 about here]

Table 4 presents models examining whether job locations explain the penalty heterogeneity by motherhood timing. Models 1-3 add firm, occupational, and industrial

[5] This penalty further decreases to just 1.7% per child if we add firm, occupational, and industrial characteristics.
[6] The fact that the pay penalties for early and average mothers are just slightly smaller than the 4.3% penalty estimated with fewer data rounds suggests that changes in the motherhood penalty as the youngest child grows up account for only a small part of the reduction in the average penalty from Models 1 to 2; most is attributable to the pay premium for late mothers. An additional analysis further shows that even with just 16 rounds of NLSY97, the motherhood penalty varies widely by age of first birth (Model 3 in Table A3 in Online Supplement).

characteristics gradually, but the penalty differences by age of first birth change only slightly. Model 4-6 remove the occupation and industry fixed effects and add the interactions between the number of children and firm characteristics, occupational skill level and broad industry in a step-by-step manner. The heterogeneity by motherhood timing remains. Taken together, these models indicate that sorting by firm type, occupation, and industry cannot explain why early and average mothers encounter a pay penalty but late mothers a premium.

[Table 4 about here]

If mothers are only disadvantaged after having a second or third child, and late mothers tend to have just one child by their late 30s, it could also explain why they encounter no penalty. Model 1 in Table 5 tests this possibility, with the results illustrated in Figure 1. The contrast is striking even when all mothers have just one child. Whereas women entering motherhood late (age 28+) generally receive a premium after childbearing,[7] women becoming mothers in their early or mid-20s experience a wage decrease with every child. Hence, late mothers' relative pay advantage cannot be explained by their number of children. Models 2-3 examine whether accounting for the youngest child's age makes a difference. Adding the presence of a very young or young child weakens the coefficients for the interaction between number of children and age of first birth, but the general pattern remains.[8] Further, the relevance of the youngest child's age does not differ by women's age of first birth. (Model 3). Model 4 shows that married mothers experience no significantly different penalty from other mothers'. Considering marital-status composition does not affect the penalty gap by motherhood timing, either.

[Figure 1 and Table 5 about here]

[7] In the figure, late mothers' pay premium appears to increase with the number of children, but the differences between the various numbers of children are not statistically significant.

[8] It is unsurprising the coefficients weaken after separately measuring the effects of having a young and a very young child, as these effects are part of the total effect of having children.

Table 6 presents fixed-effects models predicting log hourly pay by race-ethnicity.[9] Interestingly, while early Black mothers (first-birth age<22) appear to face a smaller pay penalty than their White counterparts, Black women who become mothers at an older age similarly receive a pay premium with the transition. Thus, the penalty heterogeneity by motherhood timing applies to both racial groups. The pattern for Hispanic women is less clear; regardless of the timing of first birth, this group experiences neither penalty nor premium with motherhood.

[Table 6 about here]

To assess whether women entering motherhood at various ages also change their work effort differently, Table 7 shows a fixed-effects model predicting weekly working hours (Model 1). Surprisingly, rather than altering work behavior less with the motherhood transition, late mothers' working hours decrease 6.1% *more* than early mothers. Yet the former face no greater decrease in the normative time demand at work (Model 2). Consistent with the argument that wages tend to reflect jobs' normative time demands more than workers' actual hours, Model 3 indicates no significant difference in the reduction of weekly pay with motherhood among women with differing first-birth timings. This finding contrasts sharply with the results from earlier models predicting hourly pay, which show a premium for late mothers. Altogether, the models in Table 7 indicate that the reason for late mothers' pay premium is rooted in their disproportionate reduction of work hours without a commensurate decrease in total wages. They appear to be able to adjust work hours more than earlier mothers, without being seen as less capable of meeting their occupations' normative time demand. Our additional analysis, presented in Part B in Online Supplement, indicates that this pattern cannot be explained by late mothers'

[9] We did not fit a model for those of "other" race-ethnicity because the group is small and too diverse for meaningful interpretations.

greater tendency to be in occupations that pay on a salary instead of hourly basis (e.g., professionals) or their higher likelihood to take paid maternity leave (which could reduce their average work hours). Perhaps women becoming mothers later have gained sufficient trust in their worker roles that they can avoid close supervision, making it possible not to be severely discredited when working fewer hours.

[Table 7 about here]

Conclusions

Using the most up-to-date longitudinal data, we have shown that recent mothers have become more diverse in their post-childbirth wage changes than before. Whereas for previous cohorts, women who delayed childbearing tended to face no penalty (Amuedo-Dorantes and Kimmel 2005; Taniguchi 1999), those who became mothers after their late 20s in the current cohort receive a parenthood premium comparable to men's (Killewald 2013; Yu and Hara 2021). Because of the substantial premium, adding childbirths of such mothers to the analysis leads to a modification of Yu and Kuo's (2017) conclusion that the motherhood penalty has been stable. While the magnitude of motherhood wage penalty has changed little for women having first child before their late 20s, the average penalty has declined for the new generation of mothers.

Unlike for fathers, the pay premium for late mothers does not lead to a real boost in total earnings. Because men generally keep or increase working hours with the transition to parenthood (Glauber 2008; Lundberg and Rose 2000), a fatherhood premium estimated based on hourly pay indicates a higher income. By contrast, we show that late mothers' pay premium mostly comes from their working fewer hours after childbirth without a proportional decrease in total pay. Thus, even women who delay childbearing would fall further behind men in lifetime

earnings with the transition to parenthood. Interestingly, late mothers do not experience a larger decline in the normative time demand at work than earlier mothers. This suggests that the longer time as childless workers may have earned late mothers the image of "devoted employees" (Correll et al. 2007), allowing them to comply less to their jobs' normative time demand after becoming mothers. Alternatively, late mothers may have been especially likely to work more hours than expected before childbearing, giving them room to cut back more afterwards. Either way, because employers often tie compensations to jobs' normative time demands (Yu and Kuo 2022), late mothers' comparatively large decrease in working hours does not translate into an equivalent reduction in total income.

This study contributes to the extensive literature on the motherhood pay penalty by showing the need to update our knowledge with newer data. The average motherhood penalty for the NLSY97 women is considerably lower than previous cohorts'. This trend, however, does not necessarily indicate a declining significance of motherhood penalty. Rather, with the group of women becoming increasingly heterogeneous in labor market experiences, the previous assumption of the motherhood penalty being a shared experience (albeit differences in the penalty's magnitude) for women no longer applies. At the same time, even as some women receive a motherhood premium, the implications of parenthood for their lifetime income or reasons behind their pay premium are not the same as those for men. By comprehensively examining the explanations for late mothers' pay advantage, this study sheds further light on the gendered experiences with work and family even for those faring well with parenthood.

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Table 1: Definitions and descriptive statistics for the variables

Variables	*Definition*	*Mean/ percentage*	*S.D.*
Log hourly pay^	Natural log of hourly pay (in cents), not including other nonroutine compensation, reported for the current (or latest) job at interview	6.91	0.76
Log weekly pay^	Natural log of (reported hourly pay × weekly working hours)	10.25	1.16
Log working hours^	Natural log of (total hours spent on the reported job per week)	3.36	.56
Normative time demand	Typical number of weekly hours worked by people in the occupation; created by O*NET as a continuous measure between 1 and 3, with 1, 2, 3 meaning 100% of the informants chose less than 40 hours, 40 hours, and over 40 hours per week, respectively	1.99	0.40
Number of children	Time-varying count of number of biological children alive	0.70	1.08
Age of first birth*	Age range when having first biological child[#]:		
	(1) < 22	41.91	
	(2) 22-27	26.08	
	(3) 28+	32.01	
Education	Level of education completed:		
	(1) Less than high school	17.05	
	(2) High school	18.75	
	(3) Some college	41.01	
	(4) College and more	22.74	
	(5) Unreported	0.45	
School enrollment	Currently enrolled in school	33.28	
Work experience	Total years with a job since age 14	6.88	5.11
Tenure	Total years at current job	2.18	2.79
Major employment breaks	Count of episodes of out of labor force for 6 weeks or more since first full-time job	1.95	2.29
Region	Region of residence based on Census definitions:		
	(1) Northeast	17.11	
	(2) Midwest	24.86	
	(3) South	36.76	
	(4) West	20.79	
	(5) Unreported	0.48	
Urban residence	Living in one of the three areas, based on Census definitions:		
	(1) Urban or Metropolitan Statistical area	77.29	
	(2) Rural area	22.15	
	(3) Unknown area	0.56	
Firm size	Based on the number of employees respondents reported:		
	(1) < 30	32.88	
	(2) 30-300	29.46	
	(3) >300	14.46	
	(4) Don't know	23.20	
Firm location	Measured in three categories, based on whether the firm:		
	(1) Has multiple locations	58.18	
	(2) Has a single location	29.88	
	(3) Unknown	14.46	
Firm unionization	Based on whether firm workers are:		

	(1) Unionized	8.39	
	(2) Not unionized	91.19	
	(3) Don't know	0.42	
Occupation	Measured in 3-digit codes provided by the NLSY97; 406 occupations in the sample	-	
Occupational skill level	Measured with Frederick's (2010) updated version of Hauser and Warren's (1997) indicator of occupational education for the 3-digit occupational codes	54.66	21.33
Industry	Measured in 3-digit codes provided by the NLSY97; 258 industries in the sample	-	
Broad industry	By aggregating 3-digit industry codes into 18 categories; details in Table A1 in Online Supplement	-	
Marital status	Measured in respondents' reported status at interview:		
	Never-married, not cohabiting	53.44	
	Cohabiting	16.47	
	Married	26.30	
	Divorced/separated/other	3.78	
Youngest child ages 0-3	Binary indicator, coded 1 for the youngest child present being 0-3 years old (otherwise 0)	22.67	
Youngest child ages 4-6	Binary indicator, coded 1 for the youngest child present being 4-6 years old (otherwise 0)	7.57	
Race-ethnicity*	Created by the NLSY97 based on respondents' self-reported information at first round, in 4 categories:		
	(1) Non-Hispanic White	68.07	
	(2) Non-Hispanic Black	15.20	
	(3) Hispanic	11.91	
	(4) Other	4.84	

Note: All variables without a standard deviation (S.D.) presented are categorical variables, while those with one are continuous variables, for which the mean, instead of the percentage, are shown in the column following the variable definition. The units here are person-years (N = 47,328), so the distributions are weighted by the number of years respondents appear in the specified condition or meet the criteria for the analytic sample.

^ One was added to the original number before taking the log to avoid negative or missing values.

* indicates that the variables are time-constant for each respondent; all other variables are time-varying.

[#] The distributions presented here are among those who had ever had children only (N=35,628). When we constructed the variable, we assumed the respondents who never had a child would enter motherhood two years from their last observed age, so that such respondents would remain in the models. We should note that this variable is only included as part of the interaction with the number of children. Because for those who had never been mothers, their number of children is always 0, the interaction with age of first birth is 0 regardless of which value we use for their age of first birth.

Table 2: Fixed-effects models predicting women's hourly pay

	Model 1 (Rounds 1-16)	Model 2	Model 3
Number of children	-0.043***	-0.025***	-0.034***
	(0.007)	(0.006)	(0.006)
No. of children × first birth 22-27[a]			-0.006
			(0.009)
No. of children × first birth 28+			0.076***
			(0.014)
Education (*ref.* less than high school):			
High school	-0.064***	-0.062***	-0.060***
	(0.017)	(0.016)	(0.016)
Some college	-0.016	0.004	0.005
	(0.014)	(0.014)	(0.014)
College or more	0.297***	0.342***	0.335***
	(0.021)	(0.019)	(0.019)
Unreported	0.063	0.133*	0.135*
	(0.059)	(0.061)	(0.061)
Enrolled in school	-0.087***	-0.101***	-0.100***
	(0.010)	(0.009)	(0.009)
Residential area (*ref.* rural):			
Urban	0.016	0.006	0.007
	(0.012)	(0.011)	(0.011)
Unknown	0.046	0.078	0.075
	(0.090)	(0.094)	(0.094)
Region (*ref.* Northeast):			
Midwest	-0.015	-0.029	-0.027
	(0.042)	(0.037)	(0.037)
South	-0.033	-0.023	-0.021
	(0.034)	(0.029)	(0.029)
West	0.045	0.063	0.065
	(0.044)	(0.037)	(0.037)
Unknown	-0.042	-0.094	-0.092
	(0.110)	(0.112)	(0.112)
Job tenure	0.028***	0.027***	0.027***
	(0.005)	(0.003)	(0.003)
Job tenure squared	-0.003***	-0.002***	-0.002***
	(0.001)	(0.000)	(0.000)
Work experience	0.027***	0.028***	0.025***
	(0.004)	(0.003)	(0.003)
Number of employment breaks	-0.014***	-0.013***	-0.013***
	(0.004)	(0.003)	(0.003)
Marital status (*ref.* Never-married, not cohabiting):			
Cohabiting	0.033**	0.037***	0.035***
	(0.011)	(0.011)	(0.011)
Married	0.073***	0.091***	0.083***
	(0.013)	(0.012)	(0.012)
Divorced/separated/other	0.035	0.056**	0.056**
	(0.030)	(0.021)	(0.021)
Constant	6.644***	6.660***	6.679***
	(0.045)	(0.038)	(0.038)
N (person-years)	40,108	46,997	46,997

Note: Values in parentheses are robust standard errors. The NLSY97 longitudinal weights are applied. Except for Model 1, all models use the full analytic sample. The models include individual and survey-round fixed effects.
[a] The reference group for age of first birth is younger than 22 years old.
* $p<.05$ ** $p<.01$ *** $p<.001$ (two-tailed tests)

Table 3: Characteristics of Early, Average, and Late Mothers and Nonmothers

	Age of first childbirth			No childbirth
	<22 (early mothers)	22-27 (average mothers)	28+ (late mothers)	
Hourly pay (in dollars)	14.69	18.74	30.19	20.89
	(12.93)	(13.45)	(35.15)	(19.35)
Weekly pay (in dollars)	535.28	665.11	1,162.12	807.82
	(405.21)	(506.33)	(1,434.17)	(810.12)
Weekly work hours	35.60	34.55	36.60	35.96
	(12.07)	(12.45)	(12.22)	(13.45)
Normative time demand	2.00	2.09	2.25	2.15
	(0.40)	(0.41)	(0.41)	(0.41)
Number of children	2.64	1.90	1.37	0
	(1.32)	(1.10)	(0.86)	--
Had birth after Round 16 (%)	15.64	22.94	62.75	--
Education (%):				
Less than high school	19.22	5.91	1.75	7.16
High school	32.45	21.48	7.11	11.39
Some college	39.88	41.90	25.14	36.47
College or more	7.81	30.63	65.82	44.61
Unreported	0.64	0.08	0.17	0.37
Job tenure (in years)	3.25	4.29	5.09	4.02
	(4.02)	(4.61)	(4.60)	(4.45)
Work experience (in years)	11.13	13.53	15.36	12.38
	(6.22)	(5.83)	(5.10)	(6.68)
Number of employment breaks	4.25	3.60	3.23	2.57
	(3.22)	(2.88)	(2.54)	(2.63)
Occupational skill level	50.50	60.22	71.61	62.22
	(20.84)	(22.53)	(20.74)	(22.19)
Firm size (%):				
<30 employees	31.38	31.69	27.87	29.59
30 - 299 employees	25.92	31.58	32.01	31.84
300 or more employees	15.82	16.42	24.79	20.26
Unknown	26.87	20.31	15.33	18.31
Unionization (%):				
Not in union	91.18	87.54	88.52	88.96
Unionized	7.69	11.55	11.13	9.69
Unreported	1.13	0.91	0.36	1.35
Marital status (%):				
Never married, not cohabiting	26.05	19.54	13.09	51.09
Cohabiting	19.45	13.06	11.89	18.76
Married	40.55	56.06	69.74	23.60
Divorced/separated/other	13.95	11.33	5.28	6.55
Youngest child <=3 years old (%)	27.42	27.66	52.93	--
Youngest child 4-6 years old (%)	14.82	17.76	21.46	--
Race-ethnicity (%):				
Non-Hispanic White	54.20	68.31	77.79	69.72

Non-Hispanic Black	23.94	16.23	7.79	12.62
Hispanic	17.83	12.08	8.81	10.29
Other	4.03	3.38	5.61	7.38
N (individuals)	1,172	1,154	760	1,089

Note: The characteristics are from respondents' last observed time in the analytic sample. All numbers in parentheses are standard deviations, with the numbers above the parentheses indicating mean values. The rest of the numbers are in percentage (as noted after the respective variable).

Table 4: Fixed-effects models of hourly pay, accounting for labor market locations

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Number of children	-0.035***	-0.027***	-0.029***	-0.018*	-0.009	-0.005
	(0.006)	(0.006)	(0.006)	(0.009)	(0.011)	(0.026)
# of children × first birth 22-27	0.000	0.007	0.008	0.000	0.002	0.005
	(0.009)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)
# of children × first birth 28+	0.078***	0.066***	0.071***	0.077***	0.074***	0.077***
	(0.014)	(0.013)	(0.013)	(0.014)	(0.014)	(0.014)
Firm size (*ref.* 30-299 employees):						
< 30 employees	0.003	-0.008	-0.005	0.022*	0.024*	0.017
	(0.009)	(0.008)	(0.008)	(0.010)	(0.010)	(0.010)
300+ employees	0.170***	0.090***	0.081***	0.172***	0.161***	0.118***
	(0.010)	(0.009)	(0.01)	(0.012)	(0.012)	(0.012)
Unknown	0.032**	0.014	0.016	0.040**	0.039**	0.031**
	(0.011)	(0.010)	(0.010)	(0.013)	(0.012)	(0.012)
< 30 employees × # of children				-0.029***	-0.032***	-0.022***
				(0.007)	(0.007)	(0.007)
300+ employees × # of children				-0.005	0.004	0.000
				(0.008)	(0.007)	(0.007)
Unknown × # of children				-0.011	-0.011	-0.007
				(0.009)	(0.008)	(0.008)
Firm location (*ref.* single):						
Multiple	0.04***	0.024***	0.024**	0.045***	0.048***	0.056***
	(0.008)	(0.007)	(0.008)	(0.009)	(0.009)	(0.009)
Unknown	-0.015	-0.04**	-0.041**	-0.011	-0.009	-0.007
	(0.014)	(0.013)	(0.013)	(0.017)	(0.016)	(0.016)
Multiple × # of children				-0.006	-0.006	-0.011
				(0.006)	(0.006)	(0.006)
Unknown × # of children				-0.005	-0.005	-0.011
				(0.011)	(0.011)	(0.011)
Unionization (*ref.* not in union):						
Unionized	0.125***	0.087***	0.087***	0.131***	0.131***	0.112***
	(0.01)	(0.01)	(0.01)	(0.012)	(0.012)	(0.012)
Unreported	0.000	-0.007	-0.008	-0.002	-0.002	-0.002
	(0.035)	(0.031)	(0.032)	(0.043)	(0.043)	(0.045)
Unionized × # of children				-0.008	-0.008	-0.014
				(0.007)	(0.007)	(0.007)
Unreported × # of children				0.000	0.000	-0.024
				(0.023)	(0.023)	(0.022)
Occupational skill level					0.005***	0.005***
					(0.000)	(0.000)
Occupational skill × # of children					0.000	0.000
					(0.000)	(0.000)
Human capital, marital status, & location[a]	Yes	Yes	Yes	Yes	Yes	Yes
Occupation fixed effects	No	Yes	Yes	No	No	No
Industry fixed effects	No	No	Yes	No	No	No
Broad industry dummies	No	No	No	No	No	Yes
Broad industry × # of children[b]	No	No	No	No	No	Yes
Individual fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Survey round fixed effects	Yes	Yes	Yes	Yes	Yes	Yes

Note: Values in parentheses are robust standard errors. The NLSY97 longitudinal weights are applied to all models. N = 46,997 for Models 1-4. Models 5-6 have fewer observations (N = 46,916) due to lacking skill information for some occupational codes. The reference group for age of first birth is younger than 22 years old.

[a] Same as in Model 3 in Table 2, including education, school enrollment, work experience, job tenure and tenure squared, number of employment breaks, marital status, region, and urban residence
[b] Results shown in Table A4 in Online Supplement
* $p<.05$ ** $p<.01$ *** $p<.001$ (two-tailed tests)

Table 5: Fixed-effects models of hourly pay, testing family conditions-related explanations

	Model 1	Model 2	Model 3	Model 4
Number of children		-0.036***	-0.038***	-.018*
		(0.006)	(0.007)	(0.007)
# of children × first birth 22-27		0.000	0.016	0.007
		(0.009)	(0.012)	(0.008)
# of children × first birth 28+		0.056***	0.046*	0.070***
		(0.015)	(0.024)	(0.013)
Number of children (*ref.* no child):				
1 child	-0.024			
	(0.014)			
2 children	-0.058***			
	(0.018)			
3 and more children	-0.107***			
	(0.022)			
1 child × first birth 22-27	-0.007			
	(0.022)			
1 child × first birth 28+	0.086***			
	(0.022)			
2 children × first birth 22-27	0.006			
	(0.025)			
2 children × first birth 28+	0.130***			
	(0.034)			
3 children × first birth 22-27	0.050			
	(0.034)			
3 children × first birth 28+	0.216*			
	(0.094)			
Youngest child ages 0-3		0.036**	0.043***	
		(0.011)	(0.013)	
Youngest 0-3 × first birth 22-27			-0.039	
			(0.025)	
Youngest 0-3 × first birth 28+			0.019	
			(0.034)	
Youngest child ages 4-6		0.015	0.030	
		(0.014)	(0.016)	
Youngest 0-3 × first birth 22-27			-0.058	
			(0.034)	
Youngest 0-3 × first birth 28+			-0.014	
			(0.043)	
Marital status (*ref.* never-married)				
Cohabiting	0.019	0.016	0.015	0.031**
	(0.010)	(0.010)	(0.010)	(0.012)
Married	0.052***	0.046***	0.046***	0.054***
	(0.011)	(0.011)	(0.011)	(0.013)
Divorced/separated/other	0.040*	0.037	0.038	0.056
	(0.022)	(0.020)	(0.007)	(0.032)
Cohabiting × # of children				-0.019*
				(0.008)
Married × # of children				-0.008
				(0.007)
Divorced × # of children				-0.018
				(0.014)

Note: Values in parentheses are robust standard errors. The models are estimated with the NLSY97 longitudinal

weights and include education, school enrollment, work experience, job tenure and tenure squared, number of employment breaks, marital status, region, urban residence, firm size, firm unionized, firm with multiple locations, and a series of fixed effects (occupation, industry, survey round, and individual). N= 46,997.
* p<.05 **p<.01 ***p<.001 (two-tailed tests)

Table 6: Fixed-effects models predicting women's hourly pay, by race-ethnicity

	White	Black	Hispanic
Number of children (first birth <22)	-0.043***	-0.017*	0.002
	(0.009)	(0.008)	(0.009)
# of children × first birth 22-27	0.022	0.016	-0.013
	(0.012)	(0.009)	(0.015)
# of children × first birth 28+	0.084***	0.096***	-0.008
	(0.017)	(0.027)	(0.022)
N (person-years)	23,267	12,374	9,679

Note: Values in parentheses are robust standard errors. The NLSY97 longitudinal weights are applied to all models. Each model also includes education, school enrollment, work experience, job tenure and tenure squared, number of employment breaks, marital status, region, urban residence, firm size, whether the firm is unionized, and whether the firm has multiple locations, but the coefficients are omitted to conserve space. In addition, occupation, industry, individual, and survey round fixed effects are included in all models.
† $p<.10$ * $p<.05$ ** $p<.01$ *** $p<.001$ (two-tailed tests)

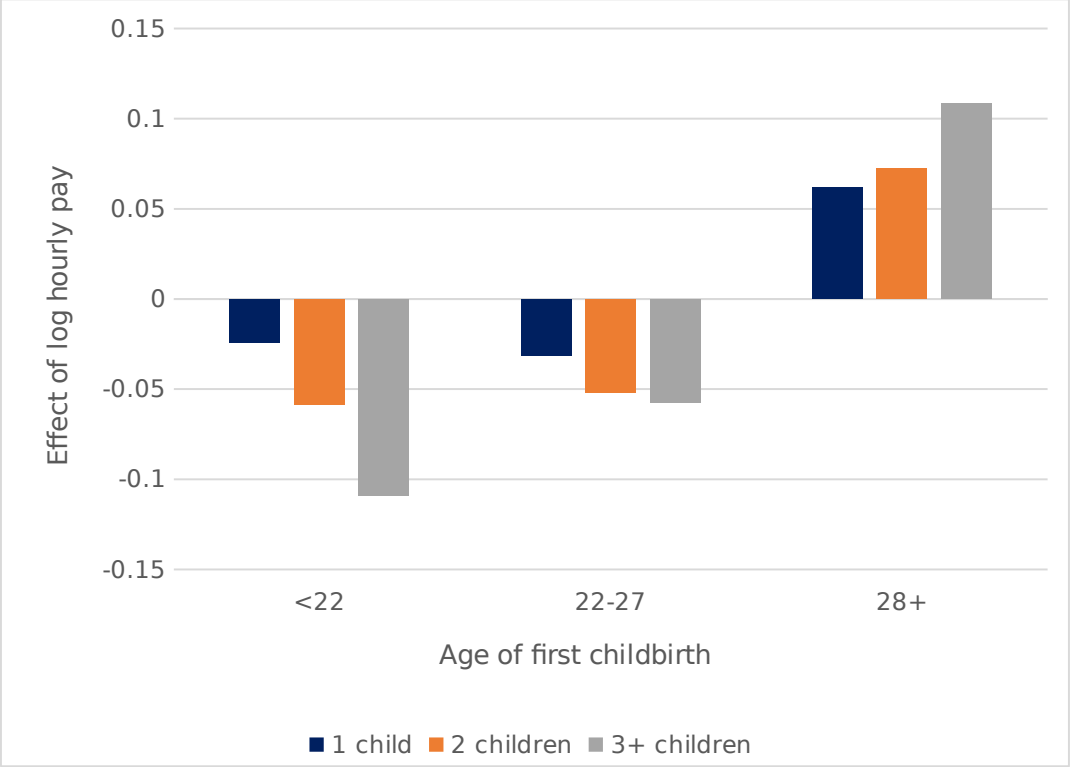
Table 7: Fixed-effects models predicting various outcomes

	Log weekly working hours	Normative time demand	Log weekly pay
Number of children (first birth. <22)	-0.038***	-0.009**	-0.074***
	(0.005)	(0.003)	(0.009)
No. of children × first birth 22-27	-0.033***	-0.010*	-0.024
	(0.008)	(0.005)	(0.013)
No. of children × first birth 28+	-0.061***	0.002	0.011
	(0.009)	(0.006)	(0.020)
Education (*ref.* less than high school):			
High school	0.148***	-0.011	0.110***
	(0.013)	(0.008)	(0.023)
Some college	0.197***	0.001	0.213***
	(0.013)	(0.008)	(0.022)
College or more	0.396***	0.198***	0.597***
	(0.018)	(0.011)	(0.031)
Unreported	0.190***	0.082**	0.234*
	(0.055)	(0.026)	(0.101)
Enrolled in school	-0.166***	-0.021***	-0.226***
	(0.007)	(0.005)	(0.014)
Residential area (*ref.* rural):			
Urban	0.012	0.000	0.013
	(0.008)	(0.005)	(0.016)
Unknown	-0.130	0.019	-0.015
	(0.121)	(0.052)	(0.162)
Region (*ref.* Northeast):			
Midwest	-0.037	0.008	-0.042
	(0.023)	(0.016)	(0.05)
South	-0.029	0.011	-0.032
	(0.019)	(0.013)	(0.038)
West	-0.029	0.028	0.05
	(0.023)	(0.016)	(0.049)
Unknown	0.258*	-0.047	0.135
	(0.130)	(0.060)	(0.177)
Job tenure	0.001	0.013***	0.021***
	(0.002)	(0.002)	(0.004)
Job tenure squared	0.000	-0.001***	-0.001***
	(0.000)	(0.000)	(0.000)
Work experience	0.017***	0.008***	0.035***
	(0.002)	(0.002)	(0.004)
Number of employment breaks	-0.013***	-0.002	-0.024***
	(0.003)	(0.002)	(0.005)
Firm size (*ref.* 30-300 employees):			
<30 employees	-0.053***	0.019***	-0.070***
	(0.007)	(0.005)	(0.013)
>300 employees	0.025***	0.013*	0.115***
	(0.008)	(0.006)	(0.015)
Not know	-0.048***	-0.004	-0.033*
	(0.008)	(0.006)	(0.015)
Firm location (*ref.* single location):			
Multiple locations	0.082***	0.020***	0.114***
	(0.006)	(0.004)	(0.012)
Not know	0.077***	0.005	0.045*

	(0.011)	(0.008)	(0.021)
Firm unionized (*ref.* no):			
Yes	0.062***	0.009	0.158***
	(0.009)	(0.007)	(0.017)
Status unknown	-0.026	-0.045	-0.051
	(0.040)	(0.026)	(0.061)
Marital status (*ref.* Never-married, not cohabiting):			
Cohabiting	0.012	-0.002	0.035*
	(0.007)	(0.005)	(0.015)
Married	-0.019*	0.008	0.033*
	(0.008)	(0.006)	(0.017)
Divorced/separated/other	0.016	0.002	0.064*
	(0.013)	(0.011)	(0.030)
Constant	3.182***	1.861***	9.833***
	(0.029)	(0.020)	(0.054)
N (person-years)	46,997	46,915	46,997

Note: The values in parentheses are robust standard errors. The NLSY97 longitudinal weights are applied when estimating the models. Both models include occupation, industry, individual, and survey round fixed effects.
* $p<.05$ ** $p<.01$ *** $p<.001$ (two-tailed tests)

Figure 1: Changes in mothers' hourly pay with the number of children, by first-birth timing



Note: Values calculated from the coefficients from Model 4 in Table 3. Symbols above or below each bar indicates whether the estimate is statistically different from 0.
† $p<.10$ * $p<.05$ (two-tailed tests)