# Rising Inequality in Mothers' Employment Statuses: The Role of Intergenerational Transmission*

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**Abstract** During the late twentieth century, U.S. mothers' propensities to hold full-time jobs became increasingly unequal across the distribution of socioeconomic status (SES). Consequently, daughters in high-SES households became likelier to be raised by working mothers that daughters in low-SES households. To what extent did this unequal exposure further shape maternal employment inequality in the twenty-first century—when these daughters had grown into adults and begun to raise their own children? Leveraging the genealogical structure of the Panel Study of Income Dynamics, this paper estimates intergenerational employment coefficients on a sample of late-twentieth century mothers and their daughters. It documents a much stronger intergenerational relationship in high-SES families than in low-SES families. Supplementary analyses find that being raised by a working mother significantly reduces the motherhood employment penalty among high-SES women, but not among low-SES women. Disparate rates of mother-daughter employment transmission by SES can account for 36% of growing inequality in maternal employment across SES groups, observed in the Current Population Survey, between 1999 and 2016. These findings indicate that intergenerational transmissions of employment advantage interact with and magnify the effects of structural forces on maternal employment inequality.

Keywords: Maternal employment, intergenerational transmission, inequality and stratification, decomposition analysis, motherhood penalty

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## Introduction

In the 1960s, few U.S. mothers were employed in paid labor; by the 2010s, college-educated mothers' employment propensities had risen dramatically, while those of mothers without college education had risen by much less. Unequal entry of mothers into the workforce is related to the widening family income distribution (Gonalons-Pons et al., 2020; Western et al., 2008), diverging investments in children (McLanahan, 2004; Guryan et al., 2008), falling intergenerational mobility (Beller, 2009), and inequality in maternal well-being (Frech and Damaske, 2012). Unequal maternal employment has become an increasingly important feature of socioeconomic inequality and consideration in the design of work-family policy (Perry-Jenkins and Gerstel, 2020).

Existing literature has studied a variety of structural factors related to mothers' employment opportunities over time and across the socioeconomic (SES) spectrum. These include changing wages (Juhn and Murphy, 1997; Leibowitz and Klerman, 1995; Pencavel, 1998); changing fertility technology and practices (Bailey, 2006; Goldin and Katz, 2002); changing affordability of child care (Attanasio et al., 2008; Averett et al., 1997; Cascio, 2009; Kubota, 2020); and the changing welfare system (Blank, 2002; Eissa and Hoynes, 2004; Hoynes and Schanzenbach, 2012). Other work has identified SES-based inequality in maternal access to work-family support—from employers (Schneider and Harknett, 2019) as well as from familial and social networks (Dow, 2016; Harknett and Hartnett, 2011). Despite considerable investigation into these factors, little attention has been paid to a factor that has arisen as their direct consequence. Specifically, as maternal employment inequality has increased, daughters in high-SES households have become likelier than daughters in low-SES households to be raised by working mothers. To what extent does this unequal exposure, via intergenerational transmission of employment propensity from mother to daughter, further employment inequality in the long run?

This paper constructs an account of maternal employment inequality that integrates structural factors, operating in society, with intergenerational channels of status attainment operating in families. Recent work has suggested that working mothers foster greater employment capacities in their daughters than do stay-at-home mothers—through the transmission of attitudes, skills and opportunities a working mother gleans from the workplace (e.g. Galassi et al., 2019; McGinn et al., 2019). Accordingly, if structural forces increase maternal employment inequality in one generation, intergenerational channels may produce a further increase in the next generation. Moreover, the structural forces that limit employment opportunities of low-SES mothers in one generation may also

limit the *rate* at which low-SES working mothers augment the employment propensities of the next generation. That is, structural forces and intergenerational channels may *interact* with one another, furthering the level of inequality beyond what would exist if this interaction were not present. These relationships have yet to be explored by the literature.

I apply this account to a quantification of rising maternal employment inequality between SES groups over the period of 1999-2016. The analysis proceeds in two steps. First, I leverage the genealogical structure of the Panel Study of Income Dynamics (PSID) to construct a data set of longitudinal employment information for over 2,200 mother-daughter pairs. The daughters were born in the mid-1960s through early 1980s and were observed as adult mothers during the twenty-first century. I use these data to estimate intergenerational maternal employment coefficients across different SES groups. Second, I feed these estimated coefficients into a decomposition framework, which I apply to population-level data from the Current Population Survey (CPS). The decomposition quantifies the contribution of disparate rates of mother-daughter employment transmission across SES groups to the evolution of SES-based inequality in maternal employment.

In addition to this primary contribution, this paper delivers three ancillary contributions. First, it updates a literature on temporal change in female employment that has primarily considered the twentieth century and analyzed employment status in the survey week (e.g. Goldin, 2006). I focus on full-time, full-year employment as a measure of labor market activity, which is more related to economic prosperity than occasional employment and has exhibited more change since 1999. Second, my usage of the PSID data improves existing intergenerational measures of maternal employment. Previous studies (e.g. Beller, 2009; Kalmijn, 1994; McGinn et al., 2019; Morrill and Morrill, 2013; van Putten et al., 2008) use the respondent's report of her mother's occupation status when the respondent was 14 or 16 years old. The longitudinal data set I construct permits a direct measurement of a mother's *entire employment history* throughout a daughter's upbringing. This yields more accurate and substantive intergenerational estimates than would be possible with a single-year measure (Solon, 1992). Third, the longitudinal data permit a supporting analysis of the effect of exposure to a working mother on one's own motherhood employment penalty. This analysis furthers the literature on determinants of motherhood penalties in the labor market.

## Maternal Employment across Time, Generations and SES Subgroups

### Measuring Change over Time

Growth in women's employment over the last 50 years has varied by demographic group and employment definition. Table 1 displays these population patterns, based on data from the March Supplement to the CPS on women aged 25-44. Both in the 1960s and the 2010s, women with pre-school-age children had lower propensities for employment than the full population of women. However, they experienced larger employment increases than the full population, especially with respect to full-time, full-year (FTFY) employment. (FTFY employment status is defined as at least 1600 hours of paid labor supplied during the year.) Moreover, although all education subgroups experienced substantial FTFY employment increases, college-educated mothers experienced the largest increase. The education differential—the ratio of the average employment propensity of college graduates to that of those with high school education or less—climbed from around one to around 1.7 between 1969 and 2016 (last row). As will be shown later, a substantial portion of this long-run increase in inequality occurred between 1999 and 2016.

In this paper, I focus on mothers with pre-school-age children, the FTFY employment concept as a measure of employment status, and the period of 1999-2016. Prior literature has primarily focused on changes in female employment during the late-twentieth century (e.g. Attanasio et al., 2008; Blau and Kahn, 2007; Cohen and Bianchi, 1999; Goldin, 2006; Juhn and Murphy, 1997; Leibowitz and Klerman, 1995; Pencavel, 1998). Though some recent accounts cover the 2000s (e.g. Abraham and Kearney, 2020; Goldin and Mitchell, 2017; Moffitt, 2012), these measure employment status during the survey week instead of employment status throughout the entire year. In addition, these studies do not directly focus on mothers. As shown in Table 1, focusing on occasional employment may understate change, especially for mothers, and especially with respect to SES-based heterogeneity. Moreover, while these studies sometimes report changes by education and demographic sub-group, they focus on aggregate change. My analysis of maternal employment throughout the early twenty-first century—with explicit attention to FTFY employment and to SES-based inequality—enriches the recent literature and provides a further update to the older literature.

### Understanding Transmission across Generations

It is reasonable to suspect that, relative to a stay-at-home mother, a working mother raises the employment status of her daughter. One theory posits that daughters adopt employment-relevant

beliefs and expectations via *socialization* by parents (Moen et al., 1997). Developmental psychologists (e.g. Eccles et al., 1990; Galambos, 2004; Jodl et al., 2001; McHale et al., 1999) have documented that children's conceptualizations of gender are affected by the behaviors of their parents. Bertrand (2019) showed that children raised by full-time-employed mothers expressed more egalitarian gender-role attitudes than children raised by comparable stay-at-home mothers. Other work has linked these attitudes to greater maternal employment (e.g. Farre and Vella, 2013; Fernandez and Fogli, 2009; Fortin, 2005). In addition, maternal employment may affect the *skills and opportunities* a daughter possesses in adulthood. Families with two working parents instead of one may earn more income, leading to greater monetary investments in their children, in turn augmenting their children's human capital (Kalmijn, 1994) and employment statuses (Beller, 2009).[1] Working mothers may also empower their daughters to demand greater work-family support from male partners and workplaces, obtain and transmit skills regarding the balance of work and family to their daughters, or secure their daughters attractive jobs through nepotistic channels (e.g. Staiger, 2020).

Drawing on increasingly available sources of intergenerational data, recent work has estimated strong intergenerational employment correlations between mothers and daughters (Galassi et al., 2019; McGinn et al., 2019; Morrill and Morrill, 2013; van Putten et al., 2008). These associations remain even after controlling for a variety of human capital and contextual factors. It is therefore plausible that some portion of long-run change in maternal employment has resulted from changing exposure of daughters to working mothers. a possibility not explored by the existing literature. This paper is one of the first, to my knowledge, to investigate this possibility—and the first to do so using intergenerationally-linked longitudinal data.[2] These data permit measurement of the entire employment history of a first-generation mother throughout the first 18 years of her daughter's life. Given the instability of maternal employment patterns over the lives of their children (Damaske and Frech, 2016; Killewald and Zhuo, 2019)—especially when children are of pre-school age—this

[1] For example, Blau and Grossberg (1992) documented a positive effect of maternal income on children's cognitive development, provided maternal employment occurs after the first year or two of the child's life.

[2] One important exception is the work of Fernandez, Fogli and Olivetti (2004) (FFO), which argued that sons exposed to working mothers grow up to be more supportive of their own female partners joining the workforce. FFO argued that part of married women's growing presence in the labor force in the late-twentieth century resulted from a growing fraction of sons being raised by working mothers. Morrill and Morrill (2013) argued that FFO's finding of a significant effect of a woman's *mother-in-law's* employment status on her own work behavior was the result of the socialization mechanism operating through the marriage market: i.e., daughters exposed to working mothers were likelier to demand egalitarian male partners. In any case, neither study utilized longitudinal data to estimate intergenerational employment coefficients and neither study addressed female employment inequality across SES groups.

measure contains substantially more information on exposure to a working mother than the single-teenage-year measures used by much of the literature.

**Interaction between Intergenerational Transmission and SES-Based Inequality**

The above literature is also limited by its focus on the average intergenerational transmission rate in the entire sample. My focus on maternal employment inequality necessitates novel computation of mother-daughter employment coefficients separately by family SES—with attention to the role of motherhood in the second (i.e. daughters') generation.[3] I develop and test the following hypotheses:

*Hypothesis 1: The intergenerational mother-daughter employment coefficient is smaller in lower-SES families.* There are three reasons to suspect this relationship. First, lower-SES women are more exposed to labor market volatility (Charles and Luoh, 2003) and are more likely to experience involuntary job loss (Kletzer, 1998). In addition, lower-SES women tend to experience less stable non-wage job features, such as unpredictable shift-scheduling and lack of job protection in case of work-family conflict (e.g. Schneider and Harknett, 2019). These forces of *employment stability* may prevent lower-SES women from consolidating the advantages of being raised by a working mother. Second, both employment as well as *family instability* may disrupt employment-relevant socialization of daughters by working mothers. Lower rates of marriage and higher rates of union disruption—disproportionately faced by low-SES mothers (e.g. Lundberg et al., 2016; Smock and Schwartz, 2020)—have been associated with less parent-child interaction and parental involvement (e.g. Kalil et al., 2014). Moreover, the greater prevalence of nonstandard work schedules among lower-SES mothers has been shown to affect their children's behavioral developments (Dunfion et al., 2013), and pit maternal well-being against quality time with children (Kalil et al., 2014). Third, ethnographic work by Lareau (2003) found that high-SES families made concerted efforts to instill individualistic ideologies and attitudes of status entitlement in their children; these efforts were less present in low-SES families. Though it is unclear whether these parenting differences are a byproduct of instability forces, the evidence indicates that the socialization channel may be weaker in lower-SES families.

*Hypothesis 2: SES-based inequality in mother-daughter employment transmission is stronger when the second (i.e. daughter) generation experiences motherhood.* There are two reasons to suspect this relationship. First, mothers who wish to maintain careers in a gendered workplace face substantial

[3] One exception is the work of Galassi et al. (2019), which includes an estimation of intergenerational employment correlations between mother and children across different maternal education statuses and family income quintiles. However, their analysis pools sons and daughters together.

work-family conflict (e.g. Damaske, 2011; Stone, 2008). The employment-relevant advantages of being raised by a working mother—through attitude, skill and opportunity formation—may particularly help an adult woman navigate the intersection of work and motherhood. Thus, high-SES mothers' employment statuses (relative to those of high-SES women) may particularly benefit from exposure to a working mother, while low-SES mothers' statuses (relative to those of low-SES women) may be particularly disrupted by employment and family instability. In addition, lower-SES mothers who wish to maintain stable employment may face greater *childcare instability*. The more unpredictable schedules of lower-SES workers induce greater reliance on informal and lower-quality sources of childcare—including no supervision at all (Harknett et al., 2020). Nonstandard work schedules themselves may impact low-SES mothers' informal support systems (Su and Dunifon, 2017). Despite the greater role of informal childcare, lower-SES households with employed mothers still spend greater income shares on childcare than higher-SES households (Blau and Currie, 2006; Laughlin, 2010). These constraints lead to greater experiences of work-family conflict and psychological distress among lower-SES mothers (e.g. Ciabattari, 2007; Dziak et al., 2010; Harknett et al., 2020). These challenges potentially disrupt employment status growth across generations of low-SES mothers.

I test this hypothesis in two ways. First, I assess how SES-based inequality in the intergenerational employment coefficient changes when I limit the second-generation sample to mothers with young children. Second, drawing on the longitudinal structure of the data, I assess SES-based inequality in how one's own motherhood employment penalty varies with one's childhood exposure to a working mother. Though it is beyond the scope of this paper to fully examine SES-based heterogeneity in the motherhood employment penalty, this hypothesis points to one additional dimension—exposure to a working mother—that may drive the motherhood penalty heterogeneity documented in previous work (e.g. Amuedo-Dorantes and Kimmel, 2005; Anderson et al., 2003; Budig and Hodges, 2010, 2014; Byker, 2016). Apart from the contemporary study of Kleven et al. (2019), which estimated intergenerational motherhood penalty coefficients using Danish administrative data, I am not aware of a study that has studied the motherhood penalty in an intergenerational framework.[4]

*Hypothesis 3: Disparate rates of mother-daughter employment transmission across SES groups explain a significant share of current maternal employment inequality. They also account for a*

[4] Moreover, Kleven et al. (2019) did not focus on SES-based inequality in the motherhood penalty, perhaps due to their relatively egalitarian Danish setting.

*significant portion of the increase in inequality during the early twenty-first century.* As shown in Table 1, SES-based inequality in maternal employment rose considerably after the 1960s. This suggests, and indeed the next section will show, that high-SES daughters growing up in the late twentieth century became likelier than low-SES daughters to be raised by working mothers. If Hypotheses 1 and 2 are correct, then this difference in *exposure* to working mothers, combined with the differential *responsiveness* of employment status to such exposure, would generate maternal employment inequality across SES groups in the twenty-first century. I test this hypothesis using a decomposition analysis that interacts population exposure to working mothers by SES with the above-specified intergenerational employment coefficients. This approach resembles that of Bloome (2017), who quantifies the portion of income inequality between individuals raised inside and outside of stable two-parent homes that can be explained by differential rates of intergenerational income transmission across childhood family structures. Both works develop the notion that structural forces of inequality, operating in society, are magnified by intergenerational processes operating in families.

## Data and Methods

### Construction and Summary of the PSID Intergenerational Dataset

The Panel Study of Income Dynamics (PSID) began following several thousand families in 1968, asking heads of households about their labor market outcomes in 1967. Using the genealogical structure of the survey, one can observe parental employment history for any individual born in 1967 or later and who belonged to one of these original families. The PSID is one of only two U.S. surveys where it is possible to observe multiple linked generations of labor force activity. The other is the National Longitudinal Survey of Youth, which contains longitudinal modules covering both an initial cohort of individuals (the NLSY79 study) and all children born to women from this initial cohort (the NLSY79 Child and Young Adult study—see Galassi et al 2019). I use the PSID instead of the NLSY as the former covers a broader range of birth cohorts. I am interested in second-generation daughters who were raising their own young children during the period of 1999-2016, and thus were typically born in the 1960s, 1970s or early 1980s. These birth cohorts are covered by the PSID, while most children of NLSY79 women were born after 1983.

I construct several summary measures of a daughter's exposure to a full-time-employed mother. Each measure takes the form *momFT_a_b*, and equals the share of years between a daughter's ages *a*

and *b* that her mother was employed full-time, full-year (FTFY). I construct *momFT_0_17*, *momFT_6_17* and *momFT_0_10*.[5] Each measure has relative strengths and weaknesses. For example, using the 6-to-17 age range results in the largest sample size, as doing so allows one to bring the 1961 through 1966 cohorts into the sample: cohorts that are left out if the 0-to-10 age range is used. On the other hand, the 0-to-10 age range is more directly related to my substantive focus, which is to explain employment behaviors of mothers in terms of their own mothers' behaviors. Thus, my main results are based on *momFT_0_10*. In any case, I verified that the 3 measures were highly correlated and produced similar results.

Next, I constructed a sample of second-generation women who are current heads of their own households or are partners of the heads. I kept only women born between 1961 and 1987—the range of years in which it is possible to reliably observe both the first-generation mother's full work history and the second-generation woman's work behavior. Because I am interested in modeling these individuals' employment behaviors relatively early in their career cycles—when they are likeliest to be raising their own children—I focus on a sample of women aged 22-49. I collected these women's mothers' person identifiers and then merged information on the women to information on their mothers' work histories using these identifiers.

Table 2 summarizes the resulting sample. Statistics are weighted by original PSID family weights. The sample contains 21,980 daughter-year observations, covering 2,228 unique daughters. The average daughter was observed around 9.9 times, with earlier birth cohorts observed more often. The average birth year of the sample is 1970, though the range of birth years covered is 1961-1987. The daughter-year observations range from the 1983 interview year to the 2015 year, with the mean observation coming from 2004. The average age of the sample is 33.5 years, roughly eight percent of the sample is black, and one-third of the sample has a four-year college degree. 44 percent of the observations correspond to "young mothers:" women who have at least one child under the age of six living in the household. These young-mother households have an average of 1.43 young children.

Table 2 also records statistics on the first-generation women: i.e. the mothers of these daughters. 41 percent of first-generation mothers had completed some college education. The average first-generation mother was employed full-time, full-year (FTFY) 27 percent of the time throughout the first 18 years of her daughter's life, with some mothers FTFY employed none of the time and some

[5] I set the variables to missing if information on the mother's work history is missing for more than half of the corresponding reference years.

mothers FTFY employed the entire time. This average propensity rises when the pre-school age range is excluded from the measure (*momFT_6_17*) and falls when the pre-school range occupies a larger share of the measure (*momFT_0_10*). Thus, there exists substantial variation across second-generation women in their mothers' employment histories.

Though my goal is to use these rich intergenerational data to inform national trends in maternal FTFY employment, it is important to note that my sample and results are only representative of the original 1968 sampling frame of the PSID (which contained relatively few Black and extremely few Hispanic families). Accordingly, I do not attempt heterogeneity analyses by race and ethnicity. In addition, I exclude Hispanic ethnicities and non-white, non-black races in the analysis of population trends in maternal employment.

**Intergenerational Regression Methodology**

*Linear Probability Models of FTFY Employment Propensity*

For a given daughter *i*, belonging to demographic sub-group *s*, observed in survey year *t*, I model her probability of working FTFY as a linear function of her mother's FTFY employment history:

$$\begin{aligned} \Pr(daughterFT_{ist} = 1) &= \alpha_{ist} + momFT_{is} \cdot \gamma_s \qquad (1) \\ &= \underbrace{\alpha_s + \mathbf{X_{it}} \cdot \beta_s}_{\text{other factors}} + \underbrace{momFT_{is} \cdot \gamma_s}_{\text{intergenerational transmission}} \end{aligned}$$

Estimating this regression model yields an estimate of $\gamma$, the intergenerational coefficient of interest. I use a linear probability model instead of a logit or a probit model for simplicity and because it is more straightforward to map a linear regression equation than a non-linear one to a decomposition analysis (e.g. Angrist and Pischke 2009). In any case, Appendix Table A1 presents logit estimates of $\gamma$ from the full sample; these estimates are almost the same as the linear estimates.

As hypothesized above, the intergenerational employment coefficient likely depends on the demographic sub-sample *s*. I present two alternative cuts of the sample on SES: one according to whether the first-generation mother had obtained any college education, and one according to whether the second-generation woman had obtained a four-year college degree.

The preferred regression specification includes controls for other economic and demographic characteristics $\mathbf{X_{it}}$. One set of "background" controls includes sets of birth cohort, parental education, own education, division-of-origin and race fixed effects. These variables control for systematic differences in women's employment status across education, race, regional and time-based

demographic groups. I also include a set of "contemporaneous" controls, designed to capture other economic and demographic determinants of women's labor supply (Killingsworth and Heckman, 1986). These include the log real hourly wage offer, partnership status, a quadratic in the male partner's earnings (set to 0 if none is present), a quadratic in the number of children aged 0-5, the number of children aged 6-10, and a quadratic in age. The real hourly wage offer is observed wage income divided by hours employed for women who were employed at all in the given reference year.[6] I impute hourly wage offers for women who were not employed at all, based on observed wages of demographically similar women, via a procedure described in the Appendix.

The interpretation of the *s*th intergenerational coefficient is the effect of a 0-to-1 change in the mother's FTFY employment history on the daughter's FTFY employment propensity, holding these background and contemporaneous variables fixed. This describes how much likelier—in percentage-point terms—an *s*-type daughter is to work full-time if her mother *always* worked full-time while raising her, versus if her mother *never* worked full-time. As my goal is to capture the extent to which a full-time-employed mother augments her daughter's capacity for full-time employment—through socialization and other unobserved investments and opportunities—I am interested in an intergenerational association between daughters' and mothers' FTFY employment propensities net of other observed determinants.

*Individual Fixed Effect Models of the Motherhood FTFY Employment Penalty*

The longitudinal structure of the data permits a parallel investigation into the intergenerational determinants of the *motherhood employment penalty*. This is the extent to which FTFY employment propensity changes when women transition into or out of raising young children. One drawback of the above analysis is that it relies on comparisons across different mother-daughter pairs, and thus does not control for systematic *unobserved* differences between families in full-time employment propensities. This analysis, on the other hand, leverages daughter fixed effect models to control for these differences. In addition to being of its own interest, this analysis can be seen as a check of the main intergenerational transmission analysis.

I begin by specifying the following regression equation:

[6] Real wage income is expressed at 2010 prices.

$$\Pr(daughterFT_{ist} = 1) = \alpha_i + \alpha_{st} + \underbrace{\mathbf{1}\{young\ kid_{it}\} \cdot \delta_{is}}_{\text{motherhood FTFY emp penalty}} \quad (2)$$

This equation contains a suite of individual fixed effects $\alpha_i$ that allow baseline full-time work propensities to vary arbitrarily across daughters. The coefficient $\delta_{is}$ describes the residual change in FTFY employment realized when daughter *i* of type *s* transitions into (or out of) raising her own young children. To distinguish motherhood effects from age effects, the $\alpha_{st}$ term contains a quadratic in the daughter's age. Thus, $\delta_{is}$ is identified from within-daughter variation in FTFY employment trajectories combined with across-daughter variation in motherhood timing.

Analogous to Eq. 1, I specify daughter *i*'s motherhood penalty, $\delta_{is}$, as a linear function of her mother's FTFY employment history:

$$\begin{aligned} \delta_{is} &= \eta_{is} + momFT_{is} \cdot \lambda_s \quad (3) \\ &= \underbrace{\epsilon_s + \mathbf{X_i} \cdot \theta_s}_{\text{other factors}} + \underbrace{momFT_{is} \cdot \lambda_s}_{\text{intergenerational transmission}} \end{aligned}$$

The intergenerational coefficient, $\lambda_s$, describes the extent to which the motherhood employment penalty faced by daughter *i* varies according to how much her own mother worked. For example, $\lambda_s = 0.1$ implies that a daughter raised by a full-time-employed mother faces a 10-percentage-point lower motherhood penalty than an otherwise-equivalent daughter raised by a stay-at-home mother. Eq. 3 also allows the motherhood penalty to vary arbitrarily according to the suite of background characteristics $\mathbf{X_i}$ introduced above: birth cohort, parental education, own education, division-of-origin and race. To recover $\lambda_s$, I plug Eq. 3 into Eq. 2 and estimate the resultant linear probability model of FTFY employment propensity.

**Measuring Population-Level Inequality in Maternal Employment: Bringing in the CPS**

To apply the above analysis to population-level trends in maternal employment, I supplement the PSID data with 1968-2017 data drawn from the Current Population Survey (CPS). The CPS is better equipped to examine population employment trends than the PSID, as it contains a bigger sample (roughly 60,000 households) and has maintained national representation over time. I use the March supplement to the CPS, which elicits annual employment information for CPS households during their

March interviews.[7] I construct a sample of women aged 22-49, white or black, and with non-missing economic and demographic information.

I use two different methods of dividing the sample by SES. One method involves grouping all women into a high- (low-) SES sub-sample who had wage offers above (below) the median of the hourly wage offer distribution, within a cell defined by survey year and potential experience bin. (Potential experience equals age minus years of education minus 6.) To construct this stratification, I adjust wage and business incomes for top-coding and compute hourly wages as total wage and business income divided by annual hours worked,[8] trimming wage outliers from the sample. I then impute hourly wage offers for non-working women via a procedure described in the Appendix.

Alternatively, I use the same definition of high SES as used in the second-generation PSID sample: having completed four years of college education. This definition has two advantages over the first one. First, it is not dependent on imputation procedures. Second, it more coherently divides the population into distinct labor markets. However, this definition has a key drawback: the underlying composition of the college and non-college populations may have changed over time as the female college attainment rate has increased. The wage-based method of stratification groups 50 percent of the population into each group in each year, and thus may be more immune to changing composition.

These stratifications reveal two important patterns of maternal employment inequality. First, Figure 1 displays widening inequality across SES groups in daughters' exposures to full-time-employed mothers.[9] (The figure uses the wage-based stratification of SES.) Consider the median-aged high-SES mother in 1999 (2016), who was born in 1966 (1983). According to Fig. 1, between the 1966 and 1983 birth cohorts, the share of high-SES daughters raised by full-time-employed mothers increased from 0.15 to 0.39—a 267 percent increase. On the other hand, the share of low-SES daughters raised by full-time-employed mothers increased far more slowly.

Second, Figure 2 and Table 3 show a large increase in maternal employment inequality between 1999 and 2016—the exact years in which the above cohorts of daughters experienced motherhood themselves. According to Table 3, the average FTFY employment-to-population ratio across the two

[7] The CPS interviews a panel of households each month for four months, then ignores the panel for eight months, and then conducts four more monthly interviews before dropping the panel from the sample.

[8] Before the 1976 survey, information used to construct annual hours worked is missing. Weeks worked information is available in the form of intervals. Using education, demographic and weeks worked interval information, I imputed annual hours in a manner described in the Appendix.

[9] The graph is based on the 0-10 daughter age range. Results are qualitatively and quantitatively similar if the 6-17 age range is used instead. Graph available upon request.

high-SES groups of mothers increased by 0.111 during this time period, or roughly 25 percent over its 1999 base rate (row 6). On the other hand, the FTFY employment-to-population ratio changed little in both groups of low-SES mothers: the average increase was only 0.001 (row 3). Thus, the *SES gap* in maternal FTFY employment rose by 11 percentage points between 1999 and 2016 (last row). By 2016, the SES gap was a full 17.6 percentage points—the highest it had ever been, at least since the 1960s when these statistics began to be tracked by the CPS.

**Decomposition Methodology**

I use Eq. 1 to analyze these SES gaps: the 17.6-percentage-point gap observed in 2016 and the 11-percentage-point growth in this gap observed between 1999 and 2016. The decomposition analyses compute the portion of the given gap attributable to intergenerational transmission of FTFY employment propensity, and the remaining portion. This remaining portion can be interpreted as structural forces—such as wages, other non-wage characteristics of available work, access to social safety nets, access to affordable childcare, and contemporaneous family structure—that shape maternal employment inequality in the absence of any intergenerational transmission mechanism. Because Eq. 1 is linear, the standard Oaxaca-Blinder decomposition methodology can be applied even though the outcome of interest—working FTFY—is binary (e.g. Angrist and Pischke 2009).

To begin, I average Eq. 1 across all daughters of type *s* observed in time window *t*. The result is the FTFY employment-to-population ratio of the given sub-group of daughters at time *t*:

$$\overline{daughterFT}_{st} = \overline{\alpha}_{st} + \overline{momFT}_{st} \cdot \gamma_s \tag{4}$$

The decomposition of the SES gap in this object straightforwardly follows:

$$\overline{daughterFT}_{HI,t} - \overline{daughterFT}_{LO,t} = \underbrace{\overline{\alpha}_{HI,t} - \overline{\alpha}_{LO,t}}_{\text{structural forces, absent intergenerational transmission}} + \underbrace{\overline{momFT}_{HI,t} \cdot \gamma_{HI} - \overline{momFT}_{LO,t} \cdot \gamma_{LO}}_{\text{structural forces, operating thru intergenerational transmission}} \tag{5}$$

For example, consider the 17.6-percentage-point gap observed in 2016. Eq. 5 decomposes this gap into the portion attributable to structural forces, operating in 2016 in the absence of any intergenerational transmission, and the portion attributable to intergenerational transmission mechanisms. The intergenerational term answers the question: "How much *lower* would the SES gap

in maternal FTFY employment be if working mothers did not directly augment their daughters' full-time employment statuses?"

The decomposition framework can also be applied to a *change* in the SES gap in daughters' FTFY employment propensities observed between two time points. This follows from the linearity of Eq. 5. For example, evaluate Eq. 5 at two different time points: $t_0$ (e.g. 1999) and $t_1$ (e.g. 2016). Let $\Delta_{t_0,t_1}(x)$ represent the increase in $x$ observed between $t_0$ and $t_1$. Then, subtracting Eq. 5 at $t_0$ from Eq. 5 at $t_1$ yields the following grand decomposition:

$$\Delta_{t_0,t_1}\left(\overline{daughterFT}_{HI} - \overline{daughterFT}_{LO}\right) = \underbrace{\Delta_{t_0,t_1}\left(\overline{\alpha}_{HI} - \overline{\alpha}_{LO}\right)}_{\text{structural forces, absent intergenerational transmission}} + \underbrace{\Delta_{t_0,t_1}\left(\overline{momFT}_{HI}\cdot\gamma_{HI} - \overline{momFT}_{LO}\cdot\gamma_{LO}\right)}_{\text{structural forces, operating thru intergenerational transmission}} \quad (6)$$

That is, Eq. 6 decomposes the observed 11-percentage-point increase in the SES gap in second-generation mothers' FTFY employment into *changes* in structural forces that favor high-SES mothers—in the absence of intergenerational transmission mechanisms—and *changes* in intergenerational transmission of employment propensities that favor high-SES mothers.

The intergenerational terms in Eqs. 5 and 6 capture the *interaction* of structural forces operating in previous generations and those operating in the current generation. To see this, consider the exposure variables $\overline{momFT}_{HI}$ and $\overline{momFT}_{LO}$. To the extent that higher-SES mothers in the previous generation encountered better employment opportunities and had more advantageous upbringings, we would expect $\overline{momFT}_{HI} > \overline{momFT}_{LO}$. Now, consider the intergenerational coefficients $\gamma_{HI}$ and $\gamma_{LO}$. To the extent that forces of instability, operating in the current generation, inhibit low-SES mothers from translating the advantages of being raised by a working mother into their own full-time employment statuses, we would expect $\gamma_{HI} > \gamma_{LO}$. The intergenerational term of Eq. 5 captures the interaction of unequal exposures to working mothers with unequal responses to being raised by a working mother. It thus describes the additional impact of structural forces on maternal employment inequality that operate intergenerationally through family channels.

As the intergenerational terms combine PSID and CPS data together, it is useful to take stock of what statistical object is computed from what dataset. Each intergenerational term calls for four numbers: second-generation mothers' average exposures to full-time-employed mothers by SES (or their changes between 1999 and 2016); and intergenerational employment coefficients by SES. Due

to its larger and more representative sample, I use the CPS to compute population-level exposure measures. (These are reported above in Fig. 1.) The intergenerational coefficients cannot be measured in the CPS. Instead, these are estimated in the PSID, as described in the preceding sub-sections. Thus, the intergenerational terms in Eqs. 5 and 6 are computed by interacting the CPS exposure measures with the PSID coefficients.

## Results

As a baseline exercise, I estimate intergenerational employment coefficients in the full PSID sample of mother-daughter pairs. I use the most comprehensive measure of the mother's employment history available—*momFT_0_17*—as my initial aim is to provide a broad description of intergenerational employment transmission that utilizes all available employment information. Table 4 reports estimated marginal effects (expressed in percentage-points /100) of a 0-to-1 change in *momFT_0_17*. Robust standard errors are clustered at the individual level and appear below the coefficients in parentheses. Appendix Table A1 presents similar estimates based on the alternative measures of the mother's employment history (*momFT_0_10* and *momFT_6_17*).

Looking at column one, a simple bivariate regression yields a positive and highly significant association between a first-generation mother's FTFY employment history and the FTFY employment behavior of her daughter. The estimate of 0.136 can be interpreted as follows: the typical daughter raised by a continuously-full-time-employed mother is 13.6 percentage points likelier to work full-time throughout her early adult life (i.e. ages 22-49) than the typical daughter raised by a stay-at-home mother. This effect amounts to 26 percent of daughters' observed FTFY employment-to-population ratio in the sample (which is 0.531). Columns 2-4 successively add in the control variables described in the preceding section. Inclusion of these controls reduces but does not largely affect the estimated intergenerational relationship: column 4, the preferred specification, reports that a 0-to-1 change in *momFT_0_17* is associated with an 11.6 percentage-point change in a second-generation mother's probability of working full-time (or 22 percent of the sample FTFY emp/pop ratio). Given the multitude of variables controlled here, the large residual association suggests that working mothers directly raise the full-time employment capacities of their daughters, through the unobserved channels described earlier (e.g. belief and attitude formation, work-family-balancing skill formation, and social

capital / nepotism). This result is quite similar in magnitude to the result of a similar analysis performed by Galassi et al. (2019) in NLSY data (see Table 5 of their study).

**SES-based Inequality in Intergenerational Employment Transmission**

Proceeding with the preferred specification, Table 5 displays heterogeneity in the intergenerational employment coefficient across motherhood and SES sub-groups. The top panel defines SES groups based on the first-generation mother's education level (no college versus some college or more), while the bottom panel stratifies on the daughter's education level (less than four years completed college versus at least four years completed college).[10] The six columns in each panel correspond to the motherhood-by-SES cuts of the sample. Due to the focus on FTFY employment in the context of raising young children, I use the 0-to-10 daughter age range to measure the first-generation mother's employment history. In doing so, I line up mother's and daughter's full-time work propensities at similar points in the life cycle (and in motherhood).

Two important patterns of heterogeneity emerge. First, the intergenerational coefficient tends to be larger in higher-SES families. Columns 1 and 2 report that a 0-to-1 increase in *momFT_0_10* is associated with a 10.2 or a 16.1 percentage-point increase in a high-SES daughter's propensity to work full-time, but only a 7.0 or 7.6 percentage-point increase in a low-SES daughter's propensity (depending on the definition of SES). Second, this SES gradient is larger in the sub-samples of young mothers. Columns 5 and 6 of Panel A report that a 0-to-1 increase in *momFT_0_10* is associated with a 16.8-percentage-point increase in a high-SES mother's full-time work propensity—but only a 3.8-percentage-point increase in a low-SES mother's full-time work propensity. Columns 5 and 6 of Panel B show a similar pattern. I cannot statistically reject equality of either pair of point estimates (*p*-values from *F* tests of equality appear below the main estimates and are in the 0.12-0.17 range), but the differences in point estimates are substantively large.

The bottom sub-panels of each panel of Table 5 record fixed effect model estimates of the motherhood FTFY employment penalty (refer to Eqs. 2 and 3). The same SES gradient shown above is evident here as well. Looking at the bottom sub-panel of Panel A, we see that a 0-to-1 increase in *momFT_0_10* is associated with a 0.142 decline in the motherhood FTFY employment penalty for high-SES daughters. That is, relative to a daughter raised by a stay-at-home mother, a daughter raised

[10] I use a more stringent definition of high SES for the second generation than for the first generation, given that college completion rates were higher in the second generation. Nonetheless, cutting the second-generation sample based on college attendance, rather than college completion, produces qualitatively similar results.

by a full-time-employed mother experiences a *14.2 percentage-point lower decline* in full-time work propensity upon raising her own young children. Given that the baseline estimate of the motherhood FTFY employment penalty in the high-SES sub-sample is -.174, this finding indicates that high-SES daughters raised by full-time-employed mothers experience small motherhood penalties. For low-SES daughters, on the other hand, a 0-to-1 increase in *momFT_0_10* is associated with only a 3.5 percentage-point decline in the motherhood FTFY employment penalty. Given a baseline motherhood FTFY employment penalty estimate of -.230, this finding indicates that a low-SES daughter experiences a large motherhood penalty regardless her own mother's work behavior. The bottom sub-panel of Panel B records a similar pattern of estimates.

These results align with the hypotheses laid out earlier. In an era of women's large-scale entry into the workforce, the mother-daughter employment coefficient reflects tension between promotive and disruptive mechanisms. For high-SES women, the intergenerational coefficient is larger in the sample of mothers than in the full sample. In addition, the motherhood employment penalty varies from small to large according to the first-generation mother's work behavior. This suggests that intergenerational promotive factors play an important role in high-SES mothers' employment outcomes. In contrast, for low-SES women, the intergenerational coefficient is small and varies little by motherhood status. In addition, the motherhood FTFY employment penalty is large and varies little according to the first-generation mother's work behavior. This suggests that for low-SES women, the employment advantages of being raised by a working mother may be mitigated substantially by disruptive forces.

**Implications for Twenty-First Inequality in Maternal Employment**

Figure 3 graphs the decompositions specified in Eqs. 5 and 6. The left bar graphs the 17.6 percentage-point gap in FTFY employment between high-SES and low-SES mothers observed in 2015-2017 CPS data. The dark portion of the bar marks the estimated share of the overall gap attributable to SES-based inequality in intergenerational factors. I compute this term in the following three steps. First, according to Fig. 1, the average exposure of high-SES daughters born in 1982-1984 to full-time-employed mothers was 0.39. (These daughters became median-aged mothers themselves in 2015-2017.) The corresponding exposure for low-SES daughters was 0.24. Second, the intergenerational coefficient for second-generation high-SES mothers is estimated at .192. (This is the average of the two estimates appearing in column 6 of Table 5). The corresponding coefficient for second-generation low-SES mothers is .062. Third, Eq. 5 yields: $0.39 \times 0.192 - 0.24 \times 0.062 =$

$0.060$. Thus, unequal intergenerational transmission of FTFY employment propensity across SES groups accounts for 6.0 percentage points of the observed 17.6 percentage-point SES gap in maternal FTFY employment. This amounts to 34.1 percent of the total gap. The whiskers on the graph trace out a 90 percent confidence interval of this estimate, following the standard error formula described by Jann (2008). The confidence interval does not overlap with zero, indicating statistical significance at the 10 percent level. (The *t*-statistic is 1.935, for a *p*-value of 0.053.)

The right bar of Figure 3 graphs an analogous decomposition of the 11.0-percentage-point increase in the maternal FTFY employment gap observed between 1998-2000 and 2015-2017. The intergenerational term is computed in the same way, except that *changes* in daughters' exposures to full-time-employed mothers, observed between the two time points, are substituted in for point-in-time exposures. As depicted in Fig. 1, high-SES daughters experienced a 24-percentage-point increase in exposure to full-time-employed mothers during the sample period, while low-SES daughters experienced a 9.4-percentage-point increase. Plugging these numbers into Eq. 6 yields: $0.24 \times 0.192 - 0.094 \times 0.062 = 0.040$. Thus, unequal *growth* in intergenerational transmission of employment propensity across SES groups accounts for an estimated 4.0 percentage points, or 36 percent, of observed *growth* in maternal employment inequality. This estimate is also statistically significant. (The *t*-statistic is 2.13, for a *p*-value of 0.033.)

## Discussion

This paper constructs an account of rising SES-based inequality in maternal employment that considers how a mother's propensity to work is conditioned by her own mother's prior work history. Before 1960, virtually no U.S. mothers worked full-time (Goldin, 2006). By 2016, over half of college-educated mothers with young children did, but they were almost 1.7 times as likely to do so as mothers with young children and only a high school education. I argue that if working mothers foster greater employment capacities in their daughters than do stay-at-home mothers, then an increase in maternal employment inequality in one generation produces a further increase in the next generation. Moreover, this effect is larger if the intergenerational transmission of employment propensity is stronger in high-SES families than in low-SES families.

To explore these hypotheses, I use the PSID to construct longitudinal employment information for over 2,200 mother-daughter pairs. I find that the decisions of high-SES women to hold on to their full-

time careers while raising young children are particularly sensitive to whether their own mothers did the same. In contrast, the intergenerational relationship is much weaker for lower-SES women and mothers. Investigation of motherhood employment penalties yields similar results: being raised by a working mother dramatically lowers the motherhood penalty faced by high-SES women, but only modestly affects the motherhood penalty faced by low-SES women. These patterns suggest that high-SES women experience strong intergenerational employment growth, as high-SES daughters effectively translate the skills and beliefs conferred by their (working) mothers into their own status attainments. In contrast, forces of instability—such as risk of job displacement; irregular work hours, unpredictable shift scheduling and employer-provided work-family support; and inconsistently affordable childcare—disrupt mother-daughter employment transmissions in lower-SES families. Applying these estimates to a decomposition of trends observed in the March CPS, I attribute 36% of the increase in maternal full-time employment inequality between 1999 and 2016 to disparate intergenerational growth in employment propensity across SES groups. Intergenerational forces have thus importantly shaped SES-based inequality in mothers' employment statuses.

By integrating intergenerational models of status attainment into the study of maternal employment, this paper enriches a literature that has mainly focused on factors such as wages, job quality, childcare availability and transfer programs. These structural forces have undoubtedly shaped maternal employment inequality (e.g. Moffitt 2012)—even in the absence of any intergenerational effects. However, this paper shows that the long-run impacts of these forces are reinforced and magnified by the presence of intergenerational transmission mechanisms. Its findings suggest that not only have high-SES mothers experienced an increase in employment opportunities relative to low-SES mothers, but that they have also become increasingly able to translate these opportunities into stable employment statuses—owing to rising structural inequality in the prior generation. The interaction of these two effects has heightened current inequality in U.S. mothers' labor market positions, compounding other dimensions of family stratification in American society.

The account of maternal employment developed in this paper reveals important implications for family-friendly workplace policy. Consider a policy that mandates all employers of low-SES service workers to provide more family-friendly work environments. These environments could feature greater leave allowances and benefits, more flexible scheduling of hours and telework allowances, or greater childcare support (Adelstein and Peters, 2019; Blau and Kahn, 2013). The traditional aim of such a policy is to reduce work-family conflict, which may induce more mothers to keep their jobs.

This paper suggests two further effects on mothers' employment propensities that operate through intergenerational channels. First, by raising the *exposure* of low-SES daughters to working mothers, the policy might further raise the employment propensity of the next generation of low-SES mothers. Second, by lessening work-family conflict and raising the stability of service-sector employment, the policy might raise the *rate* at which working mothers are able to augment their daughters' employment propensities. In sum, beyond its contemporaneous effects on the initial maternal generation, a large-scale policy mandating a family-friendlier work environment might help create a new generation of mothers more poised to take advantage of such a work environment.

These results and implications must be qualified along three important dimensions. First, although the PSID data permit precise estimation of intergenerational employment coefficients, they do not permit analysis of the mechanisms underlying these coefficients. For example, I cannot quantify the portion of the mother-daughter employment coefficient stemming from the formation of egalitarian gender-role attitudes versus unobserved opportunities. It is likely important to distinguish between these two channels, as they may have different implications for (perceived) work-family conflict faced by working mothers, and the effects of work-family conflict on maternal employment. A complete account of maternal employment inequality also requires a better understanding of how mothers socialize their daughters. Lareau's (2003) theory of concerted cultivation suggests that dual-career, high-SES families may spend considerable resources to instill gender-egalitarian and individualistic ideologies in their daughters. Such concerted efforts may not be undertaken by lower-SES families. Whether class differences in the socialization of daughters regarding employment are large and immutable, or modest and responsive to policy, are important questions.

Second, the intergenerational regression specification could be expanded to consider alternative variable definitions. This paper studied the intergenerational transmission of women's labor supply: future work may wish to consider earnings or occupational status. Here, labor supply becomes a complicating factor, as low earnings could result from a low level of hours worked or a low hourly wage. It is also difficult to assign occupation statuses to women who work zero hours, and it is unclear how to compare a mother who works few hours in a high-status occupation to a mother who works many hours in a low-status occupation. Focusing solely on labor supply skirts these issues, but at the cost of not considering these other outcomes. Moreover, the current focus on labor supply could benefit from some refinement. For example, the PSID data used here prevented a measurement of the first-generation mother's employment history beyond broad summary measures. This is because the

three measures used, based on three broadly-defined daughter age ranges (0 to 17, 6 to 17 and 0 to 10), were very highly correlated. This fact, combined with the small sample sizes, made it difficult to precisely measure the effect of maternal employment at a specific daughter age, holding maternal employment at all other daughter ages constant. Constructing a larger dataset would permit more precise intergenerational sequence analysis, which can better account for the instability of women's employment patterns over time and during the early years of their children's lives.

Third, future work should explore the intergenerational transmission of mothers' labor market outcomes in other populations. This study was based on a sample of mostly white and non-Hispanic U.S. mothers. Within the United States, sub-group analyses by race and ethnicity may shed important light on family income and labor supply differences across a diverse U.S. population. Cross-national studies should look to add depth to the broad survey of McGinn et al. (2019), which estimated intergenerational employment coefficients using cross-sectional data from 29 different countries. Longitudinal household surveys in Britain (the British Household Panel Survey) and Germany (the German Socio-Economic Panel) may permit the estimation of more detailed intergenerational employment regressions in those countries. Intergenerational population register data are also increasingly available for research in Scandinavian countries (e.g. Kleven et al., 2019): Scandinavian countries' egalitarian cultures and generous offerings of work-family supports make for potentially interesting comparisons with other Western countries.[11] Making progress along each of these three dimensions should occupy future research on maternal employment, family stratification and the efficacy of work-family policy.

[11] For early work in this direction, that used only rough one-year proxies for the first generation's employment or income, see Raaum et al (2007).

**Table 1** Employment-to-population ratios then and now, U.S. women aged 25-44

| | All women | | Mothers with young children | |
|---|---|---|---|---|
| | 1968-1970 | 2015-2017 | 1968-1970 | 2015-2017 |
| *Panel A: Shares employed > 0 hours* | | | | |
| College grads | .57 | .88 | .34 | .79 |
| Some college | .47 | .80 | .30 | .72 |
| High school or less | .46 | .66 | .28 | .59 |
| Coll/HS ratio | 1.24 | 1.38 | 1.19 | 1.34 |
| *Panel B: Shares employed > 1600 hours (full-time)* | | | | |
| College grads | .35 | .71 | .12 | .57 |
| Some college | .27 | .59 | .11 | .46 |
| High school or less | .27 | .45 | .13 | .34 |
| Coll/HS ratio | 1.32 | 1.60 | .97 | 1.69 |

*Notes:* Data source: March Current Population Survey, white or black women aged 25-44. Mothers with young children have at least one child under age six present in the household. College grads have at least four years of completed college education. Full-time, full-year (FTFY) employment is defined as supplying at least 1600 hours of paid work in the reference year.

**Table 2** PSID summary statistics

| | Mean | Std deviation |
|---|---|---|
| *Own variables (2nd generation)* | | |
| $N$ observations | 21,980 | |
| $N$ individuals | 2,228 | |
| Birth year | 1970 | 7.13 |
| Survey year | 2004 | 7.63 |
| Age | 33.54 | 7.01 |
| Black | 0.08 | 0.26 |
| 4-year college degree | 0.33 | 0.47 |
| Young child present | 0.44 | 0.50 |
| Number of young children | 0.63 | 0.82 |
| Number of young children (conditional on having one) | 1.43 | 0.62 |
| Partner's real income (1000s) | 53.76 | 51.26 |
| Partner has positive income | 0.94 | 0.23 |
| Log real hrly wage offer | 2.72 | .60 |
| *Mother's variables (1st generation)* | | |
| Some college completed | 0.41 | 0.49 |
| Share years employed FTFY... | | |
| when daughter was age 0-17 | 0.27 | 0.30 |
| when daughter was age 6-17 | 0.31 | 0.34 |
| when daughter was age 0-10 | 0.19 | 0.29 |

*Notes:* The sample consists of all PSID woman-year observations, who are members of original 1968 PSID families, meeting the following conditions: aged 22-49; born between 1961 and 1987; observed in interview years; current head of household, or spouse or unmarried partner of household head; and whose mothers had non-missing work history information. Statistics weighted by original PSID family weights. Real income and wage figures expressed in terms of 2010 US dollars. Full-time, full-year (FTFY) employment is defined as supplying at least 1600 hours of paid work in the reference year. The number of woman-year observations for momFT_0_17 and momFT_0_10 are available is 21,170.

**Table 3** Recent changes in maternal FTFY employment-to-population ratios by SES group

| | 1998-2000 | 2015-2017 | Change |
|---|---|---|---|
| Low wage | .378 | .398 | .020 |
| < 4 yrs college | .420 | .402 | −.018 |
| Low SES average | .399 | .400 | .001 |
| High wage | .490 | .587 | .097 |
| >= 4 yrs college | .440 | .565 | .125 |
| High SES average | .465 | .576 | .111 |
| **SES average gap** | .066 | .176 | .110 |

*Notes:* Data source: March Current Population Survey, white or black women aged 22-49 and with at least one young child present in the household. Full-time, full-year (FTFY) employment is defined as at least 1600 hours of paid work in the reference year. See text for description of high and low SES groups.

**Table 4** Intergenerational transmission of women's FTFY employment propensities: second generation (daughters) born in 1961-1987 and observed during ages 22-49

| | *Linear probability model estimates: 2nd gen's FTFY employment propensity* | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| 1st gen's FTFY emp history | .136*** | .123*** | .123*** | .116*** |
| | (.036) | (.037) | (.034) | (.034) |
| 2nd gen's FTFY emp/pop ratio | .531 | .530 | .503 | .500 |
| $R^2$ | .006 | .026 | .139 | .152 |
| $N$ mother-daughter pairs | 2175 | 1974 | 2154 | 1965 |
| *Controls* | | | | |
| Background controls | | ✓ | | ✓ |
| Contemporaneous controls | | | ✓ | ✓ |

*Notes:* See Table 2 for sample information. Full-time, full-year (FTFY) employment is defined as at least 1600 hours of paid work in the reference year. The explanatory variable is the share of years when the daughter was aged 0-17 that her mother worked full time. The outcome variable is a binary indicator for the daughter working full-time in adulthood. Background controls consist of fixed effects for birth cohort, race, division of origin, parental education and daughter education. Contemporaneous controls consist of the log wage offer, partnership status, a quadratic in partner earnings, a quadratic in the number of children aged 0-5, the number of children aged 6-10, and a quadratic in the individual's age. The wage offer is imputed for non-working women in a manner described in the Appendix. Robust standard errors, clustered at the individual level, appear below the coefficients in parentheses. Regressions are weighted by PSID core family sampling weights. Standard statistical significance legend used.

**Table 5** Heterogeneity in the intergenerational transmission of women's FTFY employment propensities by SES and the second generation's motherhood status

| 2nd gen's motherhood status | All | | No young kid | | Young kid | |
|---|---|---|---|---|---|---|
| SES | Low | High | Low | High | Low | High |
| **PANEL A: High SES defined as the 1st gen completing some college** | | | | | | |
| | *Linear probability model estimates:* | | | | | |
| | *2nd gen's FTFY employment propensity* | | | | | |
| 1st gen's FTFY emp history | .070 | .102* | .086* | .064 | .038 | .168** |
| | (.044) | (.058) | (.050) | (.064) | (.051) | (.068) |
| | | | | | lo=hi *p*val: .126 | |
| 2nd gen's FTFY emp/pop ratio | .495 | .506 | .588 | .617 | .373 | .394 |
| *N* mother-daughter pairs | 1116 | 796 | 892 | 666 | 850 | 595 |
| | *Fixed effect model estimates:* | | | | | |
| | *2nd gen's motherhood FTFY employment penalty* | | | | | |
| 1st gen's FTFY emp history | | | | | .035 | .142** |
| | | | | | (.058) | (.071) |
| | | | | | lo=hi *p*val: .190 | |
| 2nd gen's avg motherhood penalty, | | | | | −.230*** | −.174*** |
| 1961-1969 birth cohorts | | | | | (.017) | (.018) |
| *N* mother-daughter pairs | | | | | 1116 | 796 |
| **PANEL B: High SES defined as the 2nd gen completing 4 years of college** | | | | | | |
| | *Linear probability model estimates:* | | | | | |
| | *2nd gen's FTFY employment propensity* | | | | | |
| 1st gen's FTFY emp history | .076* | .161*** | .071 | .142** | .085* | .215*** |
| | (.041) | (.062) | (.046) | (.069) | (.048) | (.080) |
| | | | | | lo=hi *p*val: .162 | |
| 2nd gen's FTFY emp/pop ratio | .477 | .550 | .572 | .667 | .358 | .431 |
| *N* mother-daughter pairs | 1363 | 550 | 1062 | 495 | 1047 | 400 |
| | *Fixed effect model estimates:* | | | | | |
| | *2nd gen's motherhood FTFY employment penalty* | | | | | |
| 1st gen's FTFY emp history | | | | | .057 | .147* |
| | | | | | (.055) | (.082) |
| | | | | | lo=hi *p*val: .266 | |
| 2nd gen's avg motherhood penalty, | | | | | −.210*** | −.160*** |
| 1961-1969 birth cohorts | | | | | (.016) | (.020) |
| *N* mother-daughter pairs | | | | | 1363 | 550 |

*Notes:* See Table 2 for main sample information. Full-time, full-year (FTFY) employment is defined as at least 1600 hours of paid work in the reference year. Explanatory variable is the share of years when the individual was aged 0-10 that her mother worked FTFY. Linear probability model estimates come from the full specification used in Table 3. Individual fixed effect model estimates come the full specification applied to Eq. XX. Robust standard errors, clustered at the individual level, appear below the coefficients in parentheses. Regressions are weighted by PSID core family sampling weights. Standard statistical significance legend used.

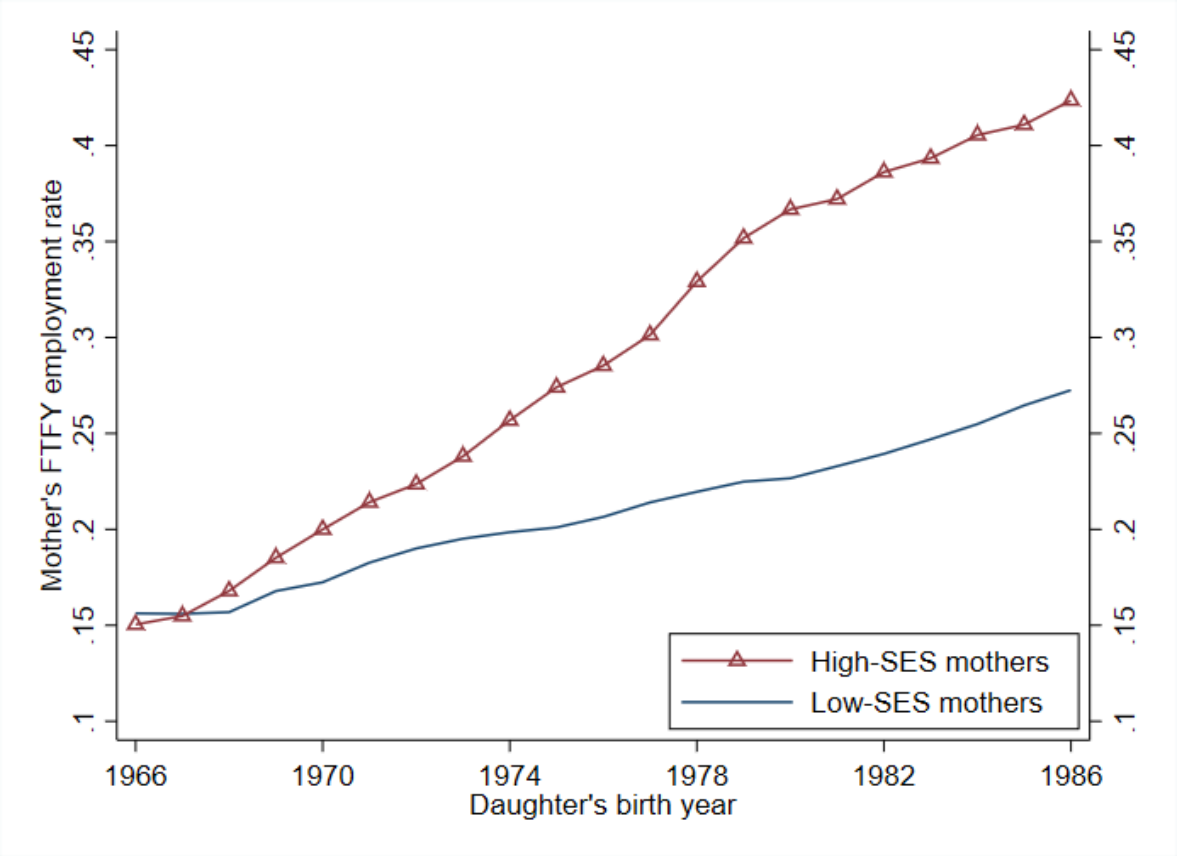


**Fig. 1** Depiction of the change in environment, from mother-at-home to mother-at-work, experienced by daughters aged 0-10 growing up in the late twentieth century: March CPS data. See text for sample description. Full-time, full-year (FTFY) employment is defined as at least 1600 hours worked in the reference year. Lines display 3-birth-cohort centered moving averages.

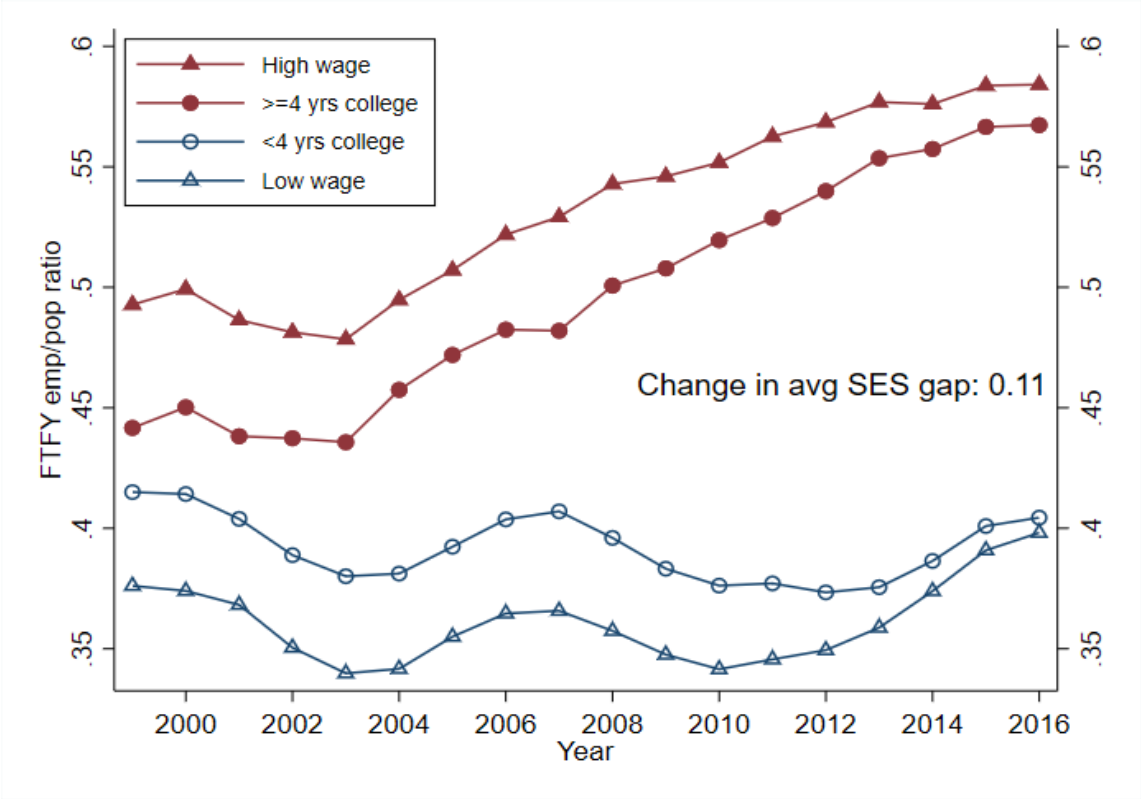


**Fig. 2** The rising SES gap in mothers' FTFY emp/pop ratios in the twenty-first century: March CPS data. Sample consists of white or black women aged 22-49 with at least one child under age 6 living at home. Lines display 3-year-centered moving averages. Full-time, full-year (FTFY) employment is defined as at least 1600 hours worked in the reference year.

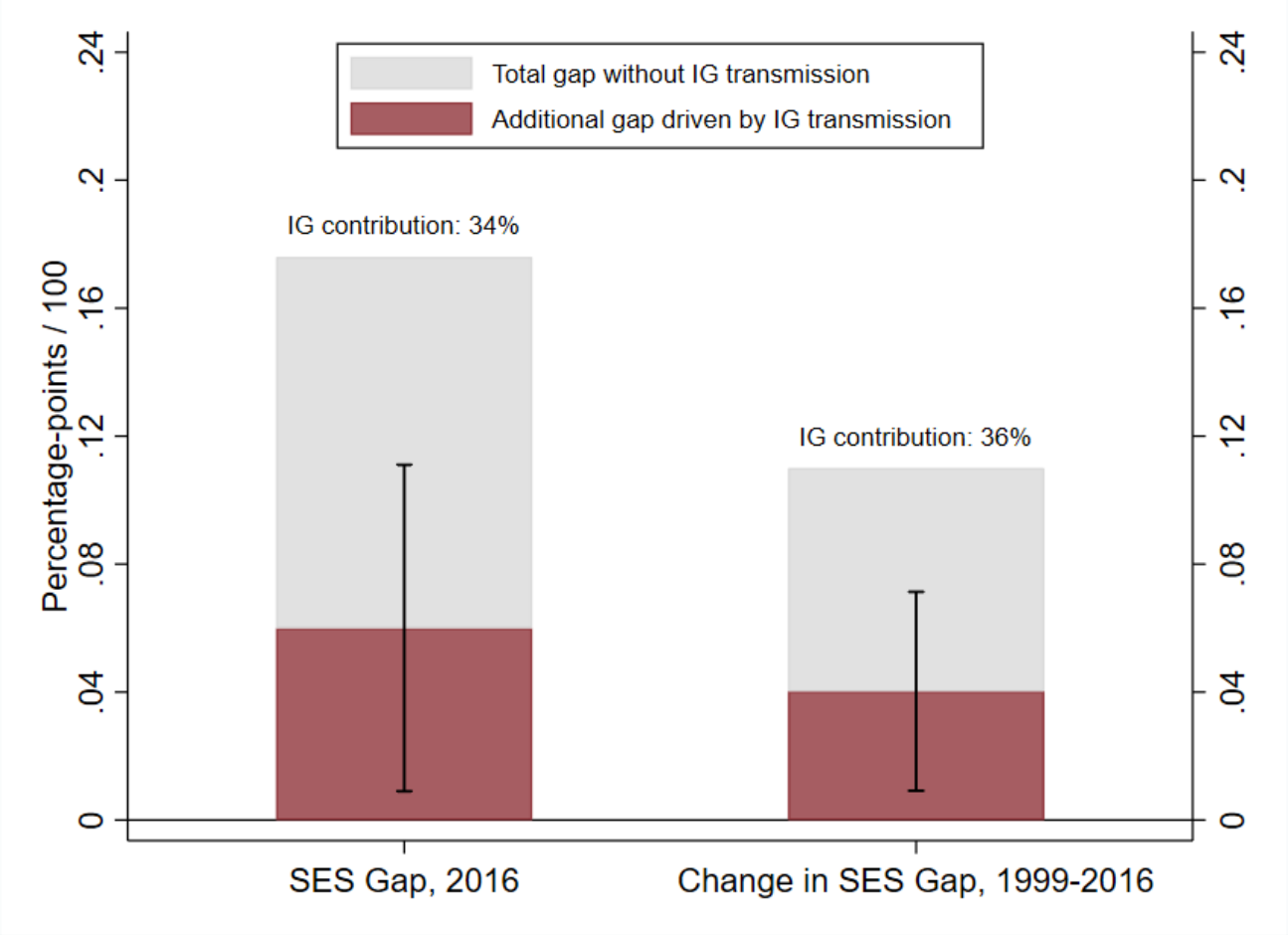


**Fig. 3** Decompositions of SES gaps in mothers' FTFY emp/pop ratios: computations based on March CPS and PSID data. Full-time, full-year (FTFY) employment is defined as at least 1600 hours worked in the reference year. The overall height of each bar comes from numbers reported in Table 3. The left bar (right bar) decomposes the observed SES gap in 2016 into the estimated intergenerational portion and a residual, as specified in Eq. 5 (Eq. 6). The whiskers trace out 90 percent confidence intervals of the intergenerational estimates.

# Appendix

**Table 6** Alternative estimates of intergenerational transmission of women's FTFY employment propensities

| | *Marginal effect estimates: 2nd gen's FTFY employment propensity* | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| **PANEL A. Linear model, independent variable: *momFT_0_10*** | | | | |
| 1st gen's FTFY emp history | .127*** | .117*** | .104*** | .100*** |
| | (.037) | (.038) | (.034) | (.036) |
| $R^2$ | .005 | .026 | .138 | .152 |
| 2nd gen's FTFY emp/pop ratio | .530 | .530 | .502 | .500 |
| $N$ mother-daughter pairs | 2097 | 1906 | 2079 | 1897 |
| **PANEL B. Linear model, independent variable: *momFT_6_17*** | | | | |
| 1st gen's FTFY emp history | .107*** | .095*** | .097*** | .095*** |
| | (.031) | (.033) | (.029) | (.030) |
| $R^2$ | .005 | .026 | .138 | .153 |
| 2nd gen's FTFY emp/pop ratio | .532 | .531 | .503 | .500 |
| $N$ mother-daughter pairs | 2225 | 2016 | 2204 | 2007 |
| **PANEL C. Logit model, independent variable: *momFT_0_17*** | | | | |
| 1st gen's FTFY emp history | .137*** | .125*** | .141*** | .133*** |
| | (.037) | (.039) | (.040) | (.041) |
| Pseudo $R^2$ | .005 | .019 | .108 | .119 |
| 2nd gen's FTFY emp/pop ratio | .531 | .530 | .503 | .500 |
| $N$ mother-daughter pairs | 2175 | 1974 | 2154 | 1965 |
| *Controls* | | | | |
| Background controls | | ✓ | | ✓ |
| Contemporaneous controls | | | ✓ | ✓ |

*Notes:* Full-time, full-year (FTFY) employment is defined as at least 1600 hours of paid work in the reference year. The explanatory variable is the share of years when the daughter was a given age that her mother worked full time. The outcome variable is a binary indicator for the daughter working full-time in adulthood. Table 4 presented estimates based on a linear model and the 0-to-17 daughter age range. This table presents estimates based on alternative age ranges (Panels A and B) or an alternative model (logit model, Panel C). See Table 4 notes for further detail on sample and regression design.

### Adjusting for top-codes in the March CPS

Around 3 percent of women in the March CPS data have top-coded earnings. The CPS separately top-codes the primary earnings source (ERN-VAL) and "other" earnings sources (WS-VAL). Before 1996, both sources of earnings were top-coded at 99,999. I replaced top-coded values with the top-code multiplied by 1.5. Between 1996 and 2010, primary earnings above the top-code threshold are

assigned the average of other top-coded earners, conditional on race, sex and full-time employment status. I did not adjust these values. Secondary earnings are top-coded at 25,000 between 1996 and 2002, and at 35,000 between 2003 and 2010. Once again I replaced top-coded secondary earnings values with the top-code multiplied by 1.5. Since 2011, and for both sources of earnings, reported earnings above the top-code threshold are systematically swapped with other values above the top-code, within a bounded interval. I elected not to adjust these values.

**Imputing annual hours worked before the 1976 March CPS**

The March CPS sample used throughout this paper begins in 1968, when children under age 15 were first tracked (so mothers with young children can be designated). Before the 1976 survey, only binned weeks worked per year and number of hours worked last week are reported, instead of the actual number of weeks worked and usual hours worked per week. For these years (1968-1975), annual hours worked were imputed in the following manner.

I began by considering a sample of women from the 1976-1981 survey years. On this sample I regressed number of weeks worked last year on a race dummy, region dummies, educational attainment dummies, a quadratic in potential experience, and education-experience interactions. I perform this regression within each weeks-worked bin (1-13, 14-26, 27-39, 40-47, 48-49, 50-52). I then imputed the number of weeks, for the 1968-1975 sample of women not in the 0-weeks-worked bin, as fitted values from these regressions. In a few cases, the fitted value was larger than 52. I replaced these fitted values with 52.

Next, I did a similar imputation for usual hours worked per week. On the 1976-1981 sample I regressed usual hours worked per week on all of the above predictors, the number of hours worked last week, the interaction between hours worked last week and the above predictors, and weeks-worked interval dummies. I imputed usual hours worked per week, for the 1968-1975 sample of women not in the 0-weeks-worked bin, as fitted values from these regressions. Imputed annual hours is the product of imputed weeks and imputed usual hours worked per week.

**Imputing wage offers for non-working women in the PSID and March CPS samples**

Hourly wages are calculated as total wage and business income divided by annual hours worked. Wages are deflated to 2010 U.S. dollars using the Consumer Price Index. To reduce measurement error, I trim calculated wages below $2.50 or above $150 from the sample.

I use the Heckman two-step procedure to impute wage offers for non-working women. First, I perform a probit regression of being employed at all in the reference year on 5 education dummies (less than high school, high school, some college, 4-year college degree, post-college), a quadratic in potential labor market experience, their interaction, education dummies interacted with a race dummy, a race dummy, education dummies interacted with region dummies, region dummies, education dummies interacted with survey year dummies, survey year dummies, marital status dummies (married, separated or divorced, never married), and marital status dummies interacted with the number of children under the age of 5.

Next, I perform a Mincerian wage regression, on the sample of working women, of log hourly wage on all of the above controls, excluding the marital status – children interaction, and including an inverse Mills ratio control function (calculated from the fitted employment probabilities estimated in the first step). Finally, I impute log wage offers for non-working women using fitted values from this regression. I iterate this imputation procedure over 6-year bins to allow the wage and selection equations to change over time. The CPS sample considers 1968-2016 survey years and the PSID

sample considers 1983-2015 survey years. Both samples consider white or black, non-Hispanic women aged 22-49.

**Constructing the high-wage-offer / low-wage-offer classification of SES in the March CPS**

The main sample consists of women aged 22-49 and with 0-35 years of potential labor market experience. I partition the main sample into cells defined by survey year and potential experience group. I consider 16 potential experience groups: 0-2, 3-4, 5-6, …, 27-28, 29-30, 31-35. Within each cell, I compute the median wage offer (i.e. the actual wage for working women or the imputed wage for non-working women) of the population-weighted sample and assign as high-SES (low-SES) each woman with a wage offer above (below) this median.

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