The Effect of State Paid Family Medical Leave Policy on Parental Childcare Time of Immigrants

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I. Introduction and Motivation

Working parents have difficulty managing work-family balance, especially regarding childcare (KPMG, 2025). Not only is it difficult to find arrangements that meet budget and time constraints, but caring for children also requires parents to respond to unexpected changes that arise during their workdays, such as sudden illness, school delays, or closures, or when previous care arrangements become unavailable. These hardships are amplified for immigrant parents: immigrants often work atypical working hours, experience more economic difficulties, face racial and social discrimination, and struggle to adjust to contrasting cultural values (Lancker, 2022). Though many foreign-born parents look to family members as an option for childcare for economic and cultural reasons, this is often no longer available to them when they immigrate to a new country (Treas & Mazumdar, 2004).

According to the National Compensation Survey, only 23 percent of civilian workers have access to employer-sponsored paid family leave (Bureau of Labor Statistics, 2023). Paid family leave is one way of alleviating the financial and time burdens working parents (and especially working immigrant parents) face. Paid Family and Medical Leave policies (PFML) refer to state-level policies that allow eligible employees to receive (partial) wage replacement and other types of compensation while taking extended time off from their job for personal medical leave, caregiving leave, and parental leave. The U.S. does not currently have a federally mandated, comprehensive PFML policy, but it does have the Family and Medical Leave Act (FMLA), which provides eligible employees with 12 weeks of unpaid, job-protected leave per year for certain family and medical reasons and the continuation of their group health benefits during the leave (U.S. Department of Labor, n.d.). Although it was aimed to ease employees' balance between work and family responsibilities, this policy alone is not sufficient in doing so, as unpaid leave is not an affordable option for many workers. As of 2025, 24 states have attempted to fill this gap for workers with their policies ranging from mandatory, social insurance policies to voluntary, private employer-based insurance plans (Bipartisan Policy, 2025).

This paper attempts to examine how state-mandated, socially insured Paid Family Medical Leave (PFML) policies across the United States affect the time spent on childcare for immigrant parents relative to non-immigrant parents. By leveraging the natural experiment created by the different times of policy adoption across states, we compare changes in time spent on childcare between immigrants and non-immigrants before and after the policies are put into effect. We use data on time spent caring for children under 18 from the 2003-2023 American Time Use Survey (ATUS) and divide our data into states by policy adoption date. We run a series of regressions measuring the effect of policy exposure and the additional effect of being an immigrant on policy exposure with varying levels of control variables. Then, we conduct two robustness checks to support the consistency of our results. Finally, we discuss the validity of the parallel trend assumption, potential bias in our models, and possible future extensions of our work, and come to the conclusion that our models cannot establish a causal relationship between policy adoption and increased caretime.

II. Literature Review

With state policy makers' growing interest in paid leave laws, there is a developing body of literature on the effect of these policies, especially on parents' outcomes. Maclean and Pabilonia, 2024

use the 2004-2022 American Time Use Survey to show that after a paid sick leave policy was adopted, parents work less and also spend more time on primary childcare. Increases in childcare time are also more pronounced for women and parents of young children. Other researchers have conducted similar quantitative studies on the effect of racial and ethnic disparities in accessing PFML (Bartel et al., 2019). indicate that Hispanic workers in general have lower access rates of PFML than their white non-Hispanic counterparts. The differences were large in raw data and remained sizable after controlling for more advanced variables indicating the sample's employment characteristics, such as industry, occupation, union membership, and part-time/full-time employment. Existing qualitative studies also show that among immigrant families, it's common for older adults, and women in particular, to travel internationally to care for (their) young grandchildren (Zhou, 2013). And among women who do provide care, recent older immigrant women provide more minutes of care than native-born and non-recent immigrant women, even after controlling for factors like ethnicity, education, and employment status (Vega, 2017). Taking on these domestic responsibilities allows for their adult children to remain in the workforce and integrate in other spheres of life in their new country more easily. Research on immigrant parents in Europe show that they're less likely to use formal childcare services than native parents, especially for recent immigrants (Lancker et al. 2019).

Our paper complements the existing research in several ways. First, we use PFML policies to more directly examine the effect on childcare time, rather than using broader paid leave policies, such as paid sick leave or paid time off, which may not always be used primarily for childcare. We also use PFML to look at the effect on parents themselves in the US, as opposed to other family members who may be providing additional or substitute care. This is especially relevant to our paper because we consider the effect on immigrants compared to non-immigrants. Immigrants often have reduced access to both formal childcare and extended family networks, which makes PFML a critical tool in enabling and affecting childcare time.

III. Experimental Design

Economic Model

To evaluate the effect of Paid Family and Medical Leave (PFML) policies on time spent on childcare, we use a differences-in-differences approach with fixed effects as it best suits our repeated cross-sectional data of observations of caretime across 50 U.S. states from 2003 to 2023. We include state fixed effects to control for state-level characteristics that differ among states, but stay constant over time, such as environmental and economic conditions, demographics, and cultural factors. We also use year fixed effects to control for time-variant differences that have an effect on childcare time across states, such as the 2020 Covid-19 Pandemic.

Our identifying assumption is the Parallel Trends Assumption, which states that in the absence of a PFML policy, PFML and non-PFML states would have experienced similar trends in caretime from 2003 to 2023. This implies that any changes in childcare time observed after a policy is adopted is from the policy, and not outside factors that would've affected childcare time anyway. Although this assumption is untestable because counterfactual outcomes are not able to be observed, it allows us to use untreated states (individuals in states that never have a PFML policy) as a control group. Due to data limitations, we use a simple difference-in-means t-test of mean childcare times in 2003 by PFML and non-PFML states to provide support for this assumption. If the differences in means is not statistically significant for each demographic group, it may strengthen the belief that there were no pre-existing differences in childcare time between states. We provide the results in Table 1.

Table 1: Difference-in-Means t-test Results (by demographic group in 2003)

Sample	t-statistic
Females	0.0646*
Males	0.5614
Immigrants	0.6894
Non-Immigrants	0.0194**
Single Parents	0.3035
Dual Parents	0.0139**
Parents with Young Children	0.5071
Parents with Non-Young Children	0.0444**

$^{*}\ p < 0.10,\ ^{**}\ p < 0.05,\ ^{***}\ p < 0.01$

All of the regressions below were run with the "robust" command to correct for heteroskedasticity and the "cluster" command to correct for serial correlation.

I. Model 1—Overall Policy Effect

$$\text{CareTime}_{it} = \boldsymbol{\beta}_0 + \boldsymbol{\beta}_1 \text{yearpolicybirth} + \boldsymbol{\alpha}_i + \boldsymbol{\gamma}_t + \boldsymbol{\varepsilon}_{it}$$

"yearpolicybirth" serves as our indicator for an individual's exposure to a state PFML policy, with "yearpolicybirth"=1 for individuals living in a state that adopted PFML before their youngest child's birth year, and 0 for otherwise. It captures the average expected effect of being exposed to a PFML policy for a given individual. Although households with children born before the PFML policy are still exposed to the policy effects, most parents primarily utilize PFML for time spent on childcare for the period during and immediately after a pregnancy. In assessing whether a state qualified as having a PFML policy, we define a state PFML policy as one that was mandatory and socially insured. The states that have a PFML policy according to this definition are as follows: California, New Jersey, Rhode Island, Washington, Massachusetts, Connecticut, Oregon, Colorado, Maryland, Delaware, Minnesota, Maine, and Washington, D.C. Our definition allows for a more accurate understanding of the effect of PFML policies because it creates the effect of a universal policy that isn't based on employer discretion. This allows us to limit the effects of possible employer-based discrimination or limits to access based on employer, so that we can get at the question of whether universally available PFML policies also translate to universally accessible ones.

II. Model 2—Additional Policy Effect on Immigrants

$$\text{CareTime}_{it} = \boldsymbol{\beta}_0 + \boldsymbol{\beta}_1 \text{yearpolicybirth} + \boldsymbol{\beta}_2 \text{immigrant} + \boldsymbol{\beta}_3 \text{immigyearpolicybirth} + \boldsymbol{\alpha}_i + \boldsymbol{\gamma}_t + \boldsymbol{\varepsilon}_{it}$$

We add an immigrant dummy variable (**$\boldsymbol{\beta}_2$immigrant**) for foreign-born individuals, including those born abroad to American parents and an interaction term (**$\boldsymbol{\beta}_3$immigyearpolicybirth**) to isolate the *additional* effect of PFML on childcare time for immigrants in PFML states relative to non-immigrants in non-PFML-implementing states. Specifically, the interaction term (**$\boldsymbol{\beta}_3$immigyearpolicybirth**) tests whether immigrants respond more strongly to PFML than non-immigrants on childcare time.

III. Model 3—Controls and Additional Policy Effect on Parents with "Young" Children

$$\text{CareTime}_{it} = \boldsymbol{\beta}_0 + \boldsymbol{\beta}_1 \text{yearpolicybirth} + \boldsymbol{\beta}_2 \text{immigrant} + \boldsymbol{\beta}_3 \text{immigyearpolicybirth} + \boldsymbol{\beta}_4 \text{singlepar} + \boldsymbol{\beta}_5 \text{immigsinglepar} + \boldsymbol{\beta}_6 \text{female} + \boldsymbol{\beta}_7 \text{immigrantfem} + \boldsymbol{\beta}_8 \text{young} + \boldsymbol{\beta}_9 \text{post\_young} + \boldsymbol{\beta}_{10} \text{hh\_numkids} + \boldsymbol{\alpha}_i + \boldsymbol{\gamma}_t + \boldsymbol{\varepsilon}_{it}$$

Model 3 adds controls for household structure (**$\boldsymbol{\beta}_4$singlepar, $\boldsymbol{\beta}_{10}$hh_numkids**), gender (**$\boldsymbol{\beta}_6$female**), and child age (**$\boldsymbol{\beta}_8$young**, a dummy variable of the youngest child aged less than 1 years old) to isolate the policy effect holding such variables constant. We also created three additional interaction terms: "immigsinglepar," "immigrantfem," and "post_young." "post_young" is an interaction term of "young" and "yearpolicybirth" that measures the expected *additional* change in time spent on childcare for parents who have a "young" child relative to non-young children, on top of the effect of being exposed to the PFML policy.

IV. Model 4— Robustness Check Through Different Policy Exposure Time

$$\text{CareTime}_{it} = \boldsymbol{\beta}_0 + \boldsymbol{\beta}_1 \text{yearpolicy} + \boldsymbol{\beta}_2 \text{immigrant} + \boldsymbol{\beta}_3 \text{immigyearpolicy} + \boldsymbol{\beta}_4 \text{singlepar} + \boldsymbol{\beta}_5 \text{immigsinglepar} + \boldsymbol{\beta}_6 \text{female} + \boldsymbol{\beta}_7 \text{immigrantfem} + \boldsymbol{\beta}_8 \text{young} + \boldsymbol{\beta}_9 \text{year\_young} + \boldsymbol{\beta}_{10} \text{hh\_numkids} + \boldsymbol{\alpha}_i + \boldsymbol{\gamma}_t + \boldsymbol{\varepsilon}_{it}$$

We then look at whether the policy effect depends on the timing of a child's birth relative to policy exposure or whether it affects time spent on childcare regardless of when the youngest child was born. We use "yearpolicy" to capture the policy effect by counting exposure through the survey year rather than the youngest child's birth year. The variable "year_young" is the interaction term of "yearpolicy" and "young" isolating the expected additional effect on childcare time for households with "young" children relative to non-young children, on top of the effect of being exposed to a PFML policy.

V. Model 5- Robustness Check Through Restricted Subsample

$$\text{CareTime}_{it} = \boldsymbol{\beta}_0 + \boldsymbol{\beta}_1 \text{yearpolicybirth} + \boldsymbol{\beta}_2 \text{immigrant} + \boldsymbol{\beta}_3 \text{immigyearpolicybirth} + \boldsymbol{\beta}_4 \text{singlepar} + \boldsymbol{\beta}_5 \text{immigsinglepar} + \boldsymbol{\beta}_6 \text{female} + \boldsymbol{\beta}_7 \text{immigrantfem} \boldsymbol{\beta}_8 \text{hh\_numkids} + \boldsymbol{\alpha}_i + \boldsymbol{\gamma}_t + \boldsymbol{\varepsilon}_{it}$$

We run a regression for parents of "young" children separately to test whether the total, not just additional, effect of the PFML policy on parents of "young" children is statistically significant. Here, "yearpolicybirth," captures the total expected policy effect and is equivalent to $\beta_1$yearpolicybirth + $\beta_9$post_young from Model 3.

The coefficient on the indicator term "yearpolicybirth" is our first coefficient of interest, as it indicates the causal effect of the PFML policy on time spent on childcare. We expect this coefficient to be positive, which is consistent with existing research suggesting that PFML policies enable parents to spend more time caring for their children with the added financial and job security. The coefficient of the interaction term "immigyearpolicybirth" captures the *additional* causal effect of PFML policy on time spent on childcare for immigrants, relative to non-immigrants. More specifically, it is the expected additional difference in childcare time for immigrants living in PFML states relative to non-immigrants, on top of the difference in childcare time for those living in non-PFML states after the policy was adopted. We expect this coefficient to be negative because of additional economic and social difficulties that immigrants face that would diminish the effect of a PFML policy, even if it still contributes to a net positive on childcare time.

Data

We use state PFML policy data from Bipartisan Policy Center, a non-profit organization that works to create bipartisan policy legislation. Information on the status of PFML policies by state was last updated on February 20, 2025, and information on state-specific legislation was drawn from corresponding state legislature websites. We measure childcare time using the variable "caretime," which is categorized on the American Time Use Survey (ATUS) as any time, in minutes, spent on physical care of a child as a primary activity in a 24 hour period. ATUS is a federally-administered, nationally representative survey that is sampled from households that participate in the Current Population Survey (CPS). ATUS data is collected through phone interviews where individuals are asked about their activities over the 24 hour period from 4 a.m. on the designated day through 3:59 a.m. on the next. We extract panel ATUS data (person/year/state) from IPUMS including the number of household children of each individual, the age of their youngest child, their marital status, their place of birth, their current state of residence, and their sex. Our analysis sample is made up of responses from 105,477 individuals (60,459 women and 45,018 men) interviewed from 2003 to 2023 who reside in all 50 states and Washington D.C.

Table A: Demographic Characteristics

Variables/Sample	N=105,477	(PFML) N=29,763 (28.22%)	(Non-PFML) N=75,714 (71.78%)
Immigrants	19,194 (18.20%)	7,849 (26.37%)	11,345 (14.98%)
Women	60,459 (57.32%)	16,700 (56.11%)	43,759 (57.8%)
Single Parents	39,281 (37.24%)	10,562 (35.49%)	28,719 (37.93%)
Female Immigrants	10,898 (10.33%)	4,443 (14.93%)	6,455 (8.53%)
Single Immigrant Parents	5,447 (5.16%)	2,168 (7.28%)	3,279 (4.33%)

Parents with Young Children	8,407 (7.97%)	2,312 (7.77%)	6,095 (8.05%)

Table A provides information on the proportion of each demographic group for all states, PFML states, and non-PFML states for the entire 20 year period. Note that immigrants appear to be overrepresented in PFML states relative to their share in the full sample. They make up 26% of the PFML-state population compared to 18% of the total. The average age of the youngest child in our dataset is 7.521 years old.

Table B: Average childcare times (2003, 2004-2023, and 2023) for all states

Year	Min	Max	Mean	Std Dev	N
2003	0	1151	66.369	104.7961	9,685
2023	0	960	77.745	118.4261	2,329

We use 2003 as the baseline pre-PFML year because it is the earliest year available in our dataset when no states adopted a PFML policy. Similarly, we use 2023 as the post-PFML year despite some states adopting a PFML policy later because it is the most recent year available in our IPUMS ATUS dataset. By 2023, eight states are counted as having adopted a PFML policy according to our definition. Table B provides summary statistics for time spent on childcare for the full sample in both 2003 and 2023, offering context for how caregiving patterns may have shifted over time. In 2003, the mean time spent on childcare was 66.3694 minutes per day, which is the average amount of time spent on at least one primary childcare activity for children ages 0 to 17. In 2023, the mean time spent on childcare was 77.7454 minutes per day.

Table C: Caretime for women by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML States	0	1151	81.378	113.1523	4,034
PFML States	0	750	87.672	116.6933	1,557

Table D: Caretime for men by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML States	0	763	43.002	85.98035	2,966
PFML States	0	700	44.735	83.43346	1,128

Table E: Caretime for immigrants by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML States	0	840	67.490	111.3021	828
PFML States	0	742	65.150	108.2564	616

Table F: Caretime for non-immigrants by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML States	0	1151	64.800	103.2783	6,172
PFML States	0	750	70.969	105.4973	2,069

Table G: Caretime for single parents by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML States	0	1151	42.932	89.04402	2,642
PFML States	0	750	39.539	83.03303	959

Table H: Caretime for dual parents by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML States	0	835	78.567	110.3377	835
PFML States	0	742	86.356	113.6574	742

Table I: Caretime for parents with young children by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML	0	1050	161.046	165.2421	570

States					
PFML States	0	715	169.797	160.2566	212

Table J: Caretime for parents with non-young children by state in 2003

Variables/Sample	Min	Max	Mean	Std Dev	N
Non-PFML States	0	1151	56.612	92.34662	6430
PFML States	0	750	61.047	95.42849	2473

Tables C-J report summary statistics for time spent on childcare for states that do and do not adopt a PFML policy by demographic group in our baseline year, 2003. We look at women and men; immigrants and non-immigrants; single parents and dual parents; and parents with “young” and non-young children in particular. Note that for both PFML and non-PFML states, women spend about twice as long on childcare than men do. Single parents and dual parents share a similar difference, but there is little difference between immigrants and non-immigrants. Parents of young children spend the longest time on childcare for both PFML and non-PFML states out of all demographic groups. Importantly, while there are baseline differences in childcare time between PFML and non-PFML states for each group, they do not, in themselves, invalidate the Parallel Trends Assumption.

IV. Results

Table 2: Regression on childcare time (minutes per day)

	(1)	(2)	(3)	(4)	(5)
	caretime	caretime	caretime	caretime	caretime
yearpolicybirth	51.88***	55.40***	35.39***		16.971
	(1.673)	(1.684)	(2.907)		(16.51)
yearpolicy				2.961**	
				(1.384)	
immigrant		2.397**	-9.120***	-8.628***	-25.315***
		(1.150)	(1.501)	(1.563)	(4.333)
immigyearpolicybirth		-8.778***	-6.500***		-3.04
		(1.254)	(1.302)		(4.835)
immigyearpolicy				-4.151**	
				(1.952)	
singlepar			-38.19***	-38.634***	-61.207***
			(1.253)	(1.556)	(4.460)
immigsinglepar			6.772**	6.561**	9.264
			(3.128)	(3.163)	(9.587)
female			39.93***	40.075***	109.04***
			(0.688)	(0.649)	(2.522)
immigrantfem			4.787**	2.048**	5.034
			(1.921)	(2.230)	(6.440)
young_child			92.59***	92.179***	
			(2.083)	(2.152)	
post_young			-18.46***		
			(4.884)		
year_young				1.927*	
				(3.202)	
hh_numkids			11.60***	11.859***	2.590
			(0.655)	(0.522)	(1.590)
_cons	67.16***	66.73***	31.33***	32.35***	111.628***
	(0.0906)	(0.233)	(0.761)	(1.60)	(3.745)
Observations	105477	105477	105477	105477	8407
R-squared	0.0104	0.0105	0.1401	0.1376	0.1393

Robust standard errors clustered by state in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Interpretation

Table 2 summarizes our main results, the expected average effect of a PFML policy on minutes spent on childcare per day. The first regression (1) does not include any controls, so it only captures the overall average effect of the policy on childcare time without accounting for any demographic differences, but still includes state and year fixed effects. As the coefficient on the "yearpolicybirth" variable shows, we find that a PFML policy leads an individual to spend 51.88 minutes more on childcare. Although all coefficients are statistically significant, the R-squared value for this model is very low, which shows that the model, in its current state, does not explain most of the variation in childcare time.

When controlling for immigrant status in our second regression (2), we find that the PFML policy effect is higher for non-immigrants than immigrants. The coefficient on the interaction term "immigyearpolicybirth" indicates that the expected additional change in childcare time after a PFML policy goes into effect is 8.77 minutes less for immigrants relative to non-immigrants, on top of the difference between PFML and non-PFML states. This difference aligns with our expectations prior to running the regressions and appears to reflect the structural or informational barriers immigrants face in accessing the benefits of the policy, despite it being a state-wide implementation. However, it is interesting to note that in the absence of a PFML policy, this model suggests that immigrants are expected to spend 2.39 minutes more on childcare than non-immigrants. Both coefficients are statistically significant, though our R-squared value only experienced a slight increase.

In our third regression (3) we separate PFML policy effects by parent structure (single and dual parents) and whether individuals have a "young" (age <1) child. With all controls in place, the PFML policy effect shown by the "yearpolicybirth" coefficient describes an expected 35.39 minute increase in childcare time for non-immigrant fathers of non-young (age>1) children in a dual-parent household after a PFML policy is adopted. This is a 12.10 minute decrease from the previous regression, and our third regression reveals that the PFML policy effect was primarily driven by parents of young (age<1) children, as they experience a 92.59 minute increase in time spent on childcare relative to parents of non-young (age>1) children. This is not surprising as parents often spend the longest time on direct care of children in this age group. The coefficient on the interaction term of "young" and the PFML policy indicator "yearpolicybirth," "post_young" reflect this relationship: parents of "young" children are expected to increase childcare time by an additional 18.46 minutes *less* than parents of older children after a PFML policy is adopted, on top of the difference between PFML and non-PFML states. This shows that while the PFML policy effect is strong overall, the relative gain is smaller for parents who already spend a significant amount of time on childcare. Contrastingly, single parents spend less than dual-parents on childcare in the absence of a PFML policy. This is not only consistent with the raw mean childcare times, but intuitively, single parents take on more responsibilities than dual parents do individually, decreasing the baseline time spent on childcare.

For immigrants, we find that the PFML policy effect, shown by the coefficient on the interaction term "immigyearpolicy,"results in an expected additional increase in childcare time by 4.151 minutes *less* than the increase experienced by non-immigrants after a PFML policy was adopted, for those living in PFML states compared to non-PFML states. We also conduct an F-test and find that the *total* effect of the PFML policy on immigrants is statistically significant, even if it is smaller in magnitude than the policy effect on non-immigrants. This model also reveals that immigrants are expected to have a lower baseline childcare time than non immigrants, holding all else constant, which differs from what we found in the second regression (2) without other demographic-level controls. All variables are statistically significant,

and the R-squared value increased considerably. We find that the coefficient estimates are less precise than previous regressions, which is a result of adding new variables, but could also be caused by irrelevant variables. However, based on the increased R-squared and adjusted R-squared values, as well as the statistically significant F-statistic and overall statistical significance of all variables, we accept that all of the variables jointly contribute to the explanatory value of our model.

Robustness Checks

We test the robustness of our main findings in two ways: using a different year for our PFML policy effect indicator (4) and using a restricted subsample (5). Although the results are not identical, we find that they support the robustness of our models and highlight the importance of how we chose to define our variable "yearpolicybirth."

First, we change how we define when an individual is counted as being exposed to a PFML policy by using the survey year instead of the birth year of their youngest child. This specification tests whether the timing of the child's birth relative to the policy is critical to the effect of the policy, or if the effect is broad enough to change parenting behavior at any year of exposure, regardless of when their youngest child was born. We find that the coefficient on "yearpolicy" decreased considerably, but is still statistically significant, showing that a PFML policy has a smaller effect on childcare time for non-immigrant fathers of non-young children who were exposed to a policy at any time compared to fathers who were exposed to a policy before their youngest child was born. Additionally, the coefficient on *"*year_young*"* is smaller in magnitude and not statistically significant, providing further evidence that a PFML policy is most effective when it is adopted before birth, rather than adopted after, even if the child is under the age of one.

Second, we restrict the subsample to individuals with "young" children (age<1) to see whether the total, not additional, policy effect on parents with "young" children is statistically significant. In this model, we find that many of our variables are no longer statistically significant, including, but not limited to "immigsinglepar," (indicator for single parents who are immigrants) and "hh_numkids" (the number of children in a household). Most importantly, however, we find that the variable "yearpolicybirth" is smaller and also no longer statistically significant. This indicates that exposure to a PFML policy does not have an effect for parents of "young" children specifically, although, based on the coefficient on "post_young" in our earlier models, the *additional* effect on caretime for "young" children relative to non-young children is statistically significant. Together, the results of our robustness checks emphasize the importance of the timing of the PFML policy compared to just the age of a child itself; the policy is more effective for parents when they're exposed to it when time spent on childcare is the longest, rather than for parents who have "young" children born prior to the PFML policy.

Potential Problems

We acknowledge potential problems in our models that may contribute to endogeneity. First, our models are subject to omitted variable bias. While we use state and year fixed effects across all of our regressions to account for both time-invariant state-specific differences and time-variant differences that affect all states, these fixed effects fail to capture differences within states that are time-variant. One such example is states passing other childcare-friendly policies while also adopting a PFML policy. If exposure to these other policies is correlated with PFML adoption and childcare time, it would bias our estimates. More specifically, it would be difficult to assume the causal effect of our model because increases in childcare time would not be attributed exclusively to PFML adoption and exposure—other policies may

also contribute. To correct for this error, we could gather more extensive data on policies that are designed to have a similar effect on parents and control for exposure using an additional variable to see whether exposure to a PFML policy holding all other policies constant still yields similar results.

There are also two potential cases of classical measurement error. First, our PFML policy indicator variable is based on the year of adoption and not the exact month of adoption, which would miscategorize some individuals. For example, "yearpolicybirth" will be 1 when a child was born in January 2003 while the state's policy went in effect 11 months later, in December 2003. This introduces random noise into our model that is correlated with neither the true error term nor the true "yearpolicybirth," so it satisfies the conditions of classical measurement error. However, because "yearpolicybirth" becomes an imperfect estimate for the true indicator, it causes endogeneity and would likely bias our estimated policy effect towards zero. Similarly, our variable for immigrant status "immigrant" is constrained by our inability to stratify immigrants by the age they immigrated due to data limitations in our IPUMS ATUS dataset. Our model may understate the policy effect, as it may be including immigrants who immigrated to the US at a younger age with their parents, allowing them to have access to family childcare arrangements that would not require them to utilize a PFML policy. We could again attempt to gather more detailed data to make our indicators have more precise categorization. We could conduct an analysis that differentiates immigrants by those who have access to family childcare and those who don't by using age or year of immigration, and see whether the effect differs for all immigrants, or only specific cohorts.

V. Conclusion

Our analysis shows that exposure to a PFML policy increases parents' time spent on childcare, while the same effect is smaller, but still statistically significant, for immigrant parents. Overall, this positive effect aligns with our hypothesis that PFML policy exposure increases time spent on childcare for non-immigrant parents but increases less so for immigrant parents. However, we conclude that the results of our study are unable to support a causal relationship between PFML policy exposure and increased childcare time due to potential cases of endogeneity, weak support for the Parallel Trends Assumption, and the low explanatory power of our models. The cases of potential omitted variable bias and measurement error in our model, if present, would cause biased estimates, making the relationship no longer causal. Secondly, while we were unable to conduct a falsification test because our dataset did not include enough years to allow us to test pre-trends, we ran a simple difference-in-means t-test for mean childcare time in 2003, the earliest pre-PFML policy year. It's important to note that our difference-in-means t-test only looks at pre-policy differences, not pre-policy trends, making it a weaker test for the Parallel Trends Assumption. As stated earlier, pre-policy differences do not necessarily denote a violation of the Parallel Trends Assumption. However, we found that differences in mean care time between PFML and non-PFML states were statistically significant for non-immigrants, dual parents, and parents of "young" children at the 5% level of significance, suggesting that for these groups, the two states may not have been comparable even before a policy was adopted. Taken together, these findings imply that there is not enough evidence to strongly believe the Parallel Trends Assumption, weakening the support for a strictly causal relationship between a PFML policy and increases in childcare time. Additionally, the highest R-squared value of all of our models is relatively low, at 0.1401. It shows that we need more independent variables to explain more variation in childcare time.

There are several ways for future research to strengthen and expand upon our findings. Future work can use additional data to conduct an event-study analysis to better strengthen our Parallel Trends

Assumption. For example, we can look at whether there was a difference in trends between PFML and non-PFML states for children who were born before policy adoption by running a regression including variables for when the youngest child's birth year is varying years earlier than the year the policy went into effect. Similarly, we can also see how the effect changes over time, such as whether there is a delayed effect, or if childcare time no longer increases years after policy exposure, by looking at the youngest children born varying years after the policy went into effect. Second, we can group children into more specific exposure categories by age, such as "never affected", "partially affected", and "always affected," to better understand how PFML impacts vary across child developmental stages. Similarly, we can group immigrants into categories by citizenship status to better understand accessibility and who actually benefits from PFML policies, even within immigrant groups. Our findings may also contribute to work on the effects of publicly-funded early childcare programs, or more generally, the effect of public investments on family well-being and outcomes.

Overall, our study shows that an increase in childcare time is positively associated with being exposed to a PFML policy by the time one's child is born and that there is a smaller increase in childcare time associated with exposure for immigrants. However, the causal effect of exposure to a PFML policy on childcare time cannot be confirmed by our models, requiring continued work in data collection and analysis.

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