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ABSTRACT

The Impact of Paid Family Leave on the Timing of Infant Vaccinations*

Raising a new-born child involves not only financial resources, but also time investment from the parents. A time constraint can affect important decisions made by parents at the early stages of an infant's life. One form of investment that is particularly important is vaccinating an infant. We analyze the impact of time constraints on immunization of infants on time. To establish a causal relationship, we exploit California's implementation of Paid Parental Leave Program as a natural experiment. Using a nationally representative dataset from the National Immunization Survey, we find evidence that the policy reduced late vaccinations for children born to parents in California after the policy was implemented. We test for heterogeneous effects of the policy on different subgroups in the population. We find the policy had a stronger impact on families that are below the poverty line. We conduct a series of falsification tests and robustness checks to test the validity of the results. In addition, our results are robust to several placebo tests.

JEL Classification: D04, I12, I18, J18

Keywords: vaccination, paid parental leave, difference-in-difference, synthetic control method

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I. Introduction

The conflict between family and work is something that many people struggle with. It might be argued that this conflict is particularly cumbersome for new parents, relative to everyone else. Given the increasing labor market participation of young women over the past few decades, the issue of paid family leave has become increasingly important. The US is the only country out of 41 OECD countries without a formal paid family leave program at the federal level (pewresearch.org). Most countries have some form of provision that allows people to take time off from work to care for one's family. These provisions can range from formal setups (paid maternity and paternity leave; child care subsidies) to informal ones (asking family members to take care of a child after birth).

In 2004, California became the first US state to implement a Paid Family Leave (PFL) policy. This policy grants individuals six weeks leave with partial wage replacement. The majority of leave takers are women and most of the leave is taken to bond with a newborn.[1] This paper analyzes the impact of California's paid family leave on infants being vaccinated on time. Vaccines are associated with several health benefits. Firstly, vaccination of individuals against certain infectious diseases like measles, tuberculosis and polio, have significant benefits for the individual receiving the vaccine. Secondly, vaccination not only protects the individual receiving the vaccine from disease, but also the population at large through herd immunity. Herd immunity results when a significant proportion of the population is vaccinated because individuals who have not received the vaccine are also protected.[2] Since children

[1] Under this policy, leave can also be taken to take care of a sick family member (spouse, father, mother, etc.).
[2] Readers can look at < https://www.vaccinestoday.eu/stories/what-is-herd-immunity/> to get a better idea of herd immunity. In this paper we do not focus on herd immunity, but take this as a given assumption. We assume that vaccinating children protects them and other children from diseases such as whooping cough, measles, etc.

under the age of six are most vulnerable to getting sick, immunization of children is particularly important from a public health standpoint. Finally, the impact of vaccines on health can have secondary positive effects on human capital formation through lower absenteeism and better performance in school (Bloom, et al. 2005). We identify the effect of paid family leave on vaccination timing by comparing children born in California before and after implementation of the policy to children born in states lacking such a policy during the same time period.

The duration of leave under California's policy is six weeks, and those taking leave can take it intermittently. Therefore, we expect that the vaccines that would be affected are those that are given to infants between the ages of one and four months. Leave taking behavior of women varies across several groups and can be influenced by factors such as mother's employment obligations, family support available to the parents, or economic conditions. While most women in the US take approximately one month leave following a birth, the policy managed to double leave taking behavior, which means that the average individual exposed to the policy would take two months instead of one month following childbirth (Rossin-Slater, et al. 2013). However, there is significant variation in leave taking among women.

To the best of our knowledge, this is the first paper to look at the effect of short-duration paid parental leave on vaccination timing in the US. We focus on timing rather than uptake because paid family leave would affect a parent's time constraint rather than attitude towards immunization. However, the timing of the vaccine can differ from child to child depending upon a parent's labor market characteristics, racial and ethnic background, health of the infant, and religious beliefs held by the family. The greater variation in vaccination timing relative to vaccination uptake helps us identify the policy's effects.

We use California's PFL mandate, which was implemented in July 1, 2004, as a natural experiment to establish a causal relationship between parental time with one's child and our outcome variables. We perform empirical analysis using the National Immunization Survey (NIS) from 2003 to 2011. We use linear probability modelling, logit regression, and synthetic control methods to test our hypothesis. Our results suggest that the policy decreased the probability of a child being vaccinated late by 5 to 6 pp after the policy was implemented. We use a series of different vaccines that are usually given to infants between the ages of one month to four months. Our results also suggest that the policy had a stronger effect on poorer families relative to everyone else.

Section II provides a summary of family leave policies in the US. We present a literature review on the topic and list how our contributions fit in Section III. Section IV provides a list of vaccines that we use as outcome variables. Section V presents the empirical model used for our analysis. We describe the data in Section VI. Section VII presents the results, along with a series of placebo tests and robustness checks. Finally, we conclude in Section VIII.

II. Family Leave in the US

The International Labor Organization (ILO) of the United Nations considers maternity leave as a fundamental human right. Most countries, if not all, have some provisions for maternity leave. There are also provisions in certain countries that provide paternity leave to ensure that new fathers have sufficient time to bond with their newborn baby. Maternity and/or paternity leave is considered important to allow new parents with the time to adjust to the role of parenting and to bond with the child. There are differences across countries regarding the duration and benefits covered by parental leave policies. Some countries are more generous (Sweden, Norway, Bulgaria for instance provides more than 40 weeks of paid leave), while others are not.[3]

Upon reviewing the parental leave policies of 185 countries and territories, the ILO found that 184 countries have some form of mandatory parental leave policies, the exception being Papua New Guinea (www.ilo.org). The first family leave legislation in the US was passed in 1993 when the Family and Medical Leave Act (FMLA) was signed into law by President Bill Clinton. According to this law, an employee working in a private or public firm with over 50 employees for 12 months is entitled to 12 weeks of **unpaid** leave, which can be used to take care of a newborn child or a sick family member. This law also protected the job of the worker who decided to go on leave. However, this law did not apply to a significant portion of the work force. According to recent data, 60 percent of workers employed in the private sector are covered by the FMLA (Klerman, et al. 2012). Furthermore, since the leave is unpaid, many low-income families voluntarily take less leave and return to work early because the foregone income can be detrimental.

During the 1960s, a few states in the US individually began implementing policies geared towards allowing pregnant women to take time off from work after childbirth. By 1969, 5 states enacted Temporary Disability Laws (TDL) to protect employees from potential income loss due to temporary disability.[4] Since pregnancy disables women temporarily, TDL provided minimal wage replacement to women after childbirth due to pregnancy. Following these five states, other states expanded existing federal regulations on leave-taking by women following childbirth. Currently, 25 states have some form of provisions that allow more generous leave benefits and/or

[3] See the following link for a detailed description of cross country analysis of paid leave: https://www.oecd.org/els/soc/PF2_1_Parental_leave_systems.pdf
[4] These states include New York, California, New Jersey, Rhode Island, and Hawaii.

durations over the federal mandates. Fourteen of these states, along with District of Columbia, reduced the firm-size requirement of at least fifty employees to as low as ten. The remaining states increased the length of the leave. It is important to note that beside the five states with TDL, the rest do not compensate for the lost income.[5]

In 2002, California became the first state to introduce a state sponsored family leave law, which provides leave-takers with approximately 55% of their forgone wages with a cap of $1,216 per week.[6] This law, implemented in July 2004, allowed all private sector employees irrespective of firm size up to six weeks of paid leave to take care of a newborn, or a newly adopted child, or a sick family member. Following California, Washington, New Jersey, Rhode Island, and New York passed their own legislation to have paid family leave mandates like that of California in 2007, 2008, 2013, and 2016 respectively. Washington has yet to implement the policy due to budget issues, while the other states have already either implemented the policy or will do so in the future (New York started phasing in the benefits of the policy beginning in 2018).

PFL in California, New Jersey and Rhode Island are very similar in structure. California and New Jersey provide six weeks leave while Rhode Island provides 4 weeks. All three states provide partial wage replacement: California (55%), New Jersey (66%) and Rhode Island (60%). The program is funded through employee payroll taxes. The tax rate varies across the three states: California (0.9% of the first $106,742), New Jersey (0.08% on the first $32,000), and Rhode Island (1.2% on the first $64,200). The benefits are capped to a maximum amount, which also varies across the three states.[7] All three states have implemented paid leave through their TDI insurance

[5] http://www.nationalpartnership.org/research-library/work-family/expecting-better.pdf
[6] This is the cap as of January 1st, 2018. Also, beginning January 1st, 2018, individuals can receive 60 to 70 percent of wage replacement.
[7] Currently California's maximum wage replacement cap is $1216/week, New Jersey's is $615/week, and Rhode Island's is $817/week.

framework. Since its implementation, California's PFL covered 18 million people, with 1.9 million claims being filed by parents to bond with newborn children.[8]

New York's PFL policy is the most generous one. Passed in 2016, the phase-in is from 2018 through 2021. Beginning January 2018, eligible employees received eight weeks of leave with 50 percent wage replacement. By 2021, the policy will entitle employees, who have worked full-time for 26 weeks or part-time for 175 days, 12 weeks of paid leave with 67 percent wage replacement. Like the programs in California, New Jersey and Rhode Island, New York's PFL will be funded by employees. Payroll deductions started in July 2017.[9]

Currently, there is no US federal law regarding paid parental leave in a form similar to what California, New Jersey, and Rhode Island have in place. However, there are on-going discussions on both the Democratic side as well as the Republican side of the political spectrum. With a changing labor market, and more women receiving education in technical fields, people have begun to promote federal parental leave legislation.

III. Literature Review

Existing papers focus on how maternal employment affects vaccination take-up rates and timing. Berger, et al. (2005) use longitudinal data from National Longitudinal Survey of Youth (NLSY) to see how early maternal employment affects breastfeeding rates and immunizations. Using OLS and propensity score matching techniques they find that women who return to work before 12 weeks following childbirth are associated with lower on-time immunizations. Our paper differs from theirs in the empirical approach. We use California's Paid Family Leave (PFL)

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[8] http://www.nationalpartnership.org/research-library/work-family/paid-leave/paid-leave-works-in-california-new-jersey-and-rhode-island.pdf
[9] https://www.ny.gov/programs/new-york-state-paid-family-leave

insurance program as a natural experiment and employ a difference-in-difference estimation technique to establish causality.

Fatiregun, et al. (2012) found maternal employment to be a determining factor responsible for low vaccination rates among children aged 12-23 months in a southern district of Nigeria. In contrast, Mindlin, et al (2009) performed a systematic review of child health in pre-school children from OECD countries and found that children born to employed mothers have just as good or better vaccination uptakes compared to those born to unemployed mothers. Their results might be driven by other factors such as greater income that becomes available to families when the mother is working. Therefore, it might be important to factor in the context in which the study is conducted to account for infrastructure or economic differences.

Hajizadeh et al. (2015) use longitudinal evidence from 20 low and middle-income countries and find that more generous paid maternity leave mandates had a positive effect on vaccination rates for DTP (Diphtheria, Tetanus, and Pertusis) but no effect on BCG or Polio vaccination rates. Their paper supports the view that offering generous paid maternity leave can improve vaccination rates. Similar results are found by Daku, et al. (2012) when they analyze the effect of maternity leave provisions on vaccination coverage rates. Another paper by Udea et al. (2014) uses logistic regression on survey responses from Japan, and finds that when Japanese women take parental leave, their children were significantly less likely to be behind schedule on vaccination.

However, these papers consider countries that might have very different labor market prospects for women and men than the US. Also, the institutions and attitudes towards family in these countries might be different from that in the US. As such, our paper studies how paid parental leave affects vaccination timing for infants in the US context.

IV. Vaccines Used in the Paper

To conduct our analysis, we use *Hepatitis-B* (HepB), *Diphtheria Tetatus Pertusis* (DTP), *Haemophilus Influenza* Type *B* (HIB) vaccines. The reason behind using these vaccines is that they are administered in multiple doses. The first dose of the three vaccines is given when the infant is born, and the follow up doses are required to be given before he/she is six months old. Since the leave is exhausted by the time the infant is six months, these vaccines are perfect to test our hypothesis that parental time is a significant constraint causing delay in vaccination.

Hepatitis B is a serious disease that is caused by the *Hepatitis B Virus* (HBV). It usually attacks the liver and can cause chronic health issues such as cirrhosis of the liver, liver cancer, liver failure, and even death. The danger behind contracting Hepatitis B is that chronically sick people may not be aware of their infection and might pass the virus onto other unsuspecting people. According to CDC, approximately 1.4 million people in the United States might be chronically infected with HBV. Approximately 90% of infants who get infected will become chronically ill, resulting in one out of four deaths.[10] Since the second dose of *HepB* is given when the infant is between one and two months, it serves as the best vaccine to test our hypothesis. Given Rosin-Slater, et al.'s (2012) results, the individual exposed to the policy would increase leave taking from one month to two months. This increased leave-taking should lead to fewer infants being vaccinated late with the second dose of *HepB* vaccine. This is true if time and financial constraints facing new parents are some of the reasons behind delaying immunization.

DTP is a popular vaccine designed to protect individuals against three types of harmful diseases: Diphtheria, Tetanus, and Pertussis. Caused by *Corynebacterium diphtheriae*, diphtheria

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[10] https://www.cdc.gov/vaccines/hcp/vis/vis-statements/hep-b.html

results in a thick covering in the back of the throat and may lead to breathing difficulties, heart failure, paralysis, and even death.[11] Tetanus is caused by the *Clostridium tetani* bacteria and can cause painful muscle contractions. It is also commonly called lockjaw and can create difficulties in swallowing.[12] Pertussis or whooping cough is a highly contagious disease caused by bacterium *Bordetella pertussis*, and can cause uncontrollable, violent coughs that make breathing difficult.[13] The incidence of pertussis has been steadily increasing in the United States according to recent reports from CDC.

The third vaccine used in this paper is the *Haemophilus Influenza* Type *B* (*HIB*) vaccine. *HIB* is a very serious disease that usually affects children less than five years old. It is caused by the *Haemophilus Influenzae* bacterium. When infected, it can lead to pneumonia, meningitis, bacteremia, and epiglottitis. One in five children surviving the infection will develop brain damage and/or hearing loss. The availability of the *HIB* vaccine has drastically reduced the number of serious cases of infection since 1991.[14]

V. Empirical Specification

Our main objective is to determine whether parents are delaying immunization of infants because of time constraints they face due to work obligations. To test this directly, we use a difference-in-difference framework, which compares the outcome for our treated group (children

[11] https://www.cdc.gov/diphtheria/index.html
[12] https://www.cdc.gov/tetanus/index.html
[13] https://www.cdc.gov/pertussis/index.html
[14] https://www.vaccines.gov/diseases/hib/index.html

born in California after the policy implementation in July 1st, 2004) to a control group (children born in states that lack paid parental leave).[15] We estimate two models:

$$y_{ist} = \beta_0 + \boldsymbol{\beta_1}(Treat * Post)_{ist} + \delta_s + \delta_t + \varepsilon_{ist} \qquad \text{Model (1)}$$

$$y_{ist} = \beta_0 + \beta_1(Treat * Post)_{ist} + \boldsymbol{\beta_2}(Treat * Post * Poor)_{ist} + \boldsymbol{X}'\boldsymbol{A} + \delta_s + \delta_t + \varepsilon_{ist} \qquad \text{Model (2)}$$

The first specification, Model 1, is a simple two-way fixed effects model, with state and year fixed effects. This model controls for time-invariant but state-specific factors, and state-invariant but time specific factors that might affect the outcome variables of interest. The coefficient of ***Treat*Post*** is the main coefficient of interest in Model (1) and simply gives the average effect of the policy on the entire population.

Our second specification, Model 2, exploits an income-to-poverty ratio variable which indicates whether the selected child is born to a family that is below the poverty threshold conditional on family size and family income. The Census Bureau computes an income threshold level using information on the family size, age of family members, and earnings of the members to determine poverty thresholds. If the family income is below the threshold then that family, and everyone in the family, is considered to be living under the poverty line. This threshold varies by family composition (size and age) and is also updated every year to adjust for inflation.[16]

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[15] We drop District of Columbia and New Jersey from our analysis. Since District of Columbia has paid parental leave that is funded by employers instead of employees, including DC in the control group would be wrong. New Jersey implemented a similar type of policy in 2009.

[16] For instance, if income-to-poverty ratio is 1.25, then the family is 25 percent above the poverty threshold. If the ratio is 0.75, then the family is 25 percent below the poverty threshold.

We construct a dummy variable ***Poor***, which equals 1 if the selected child is from a family that is below this poverty threshold. Given that previous studies found the policy to benefit less-advantaged groups more, we expect that poorer families will be impacted more relative to non-poor families. This might also be true because poorer families are more likely to be working in jobs that are non-unionized and lack formal access to paid parental leave.[17]

Our outcome variable of interest, $\boldsymbol{y_{ist}}$, equaling 1 if the selected infant was vaccinated after the recommended schedule (implying late vaccination), and 0 if vaccinated on time. Coefficients δ_s, δ_t represent state and year fixed effects. State fixed effects capture state specific but time-invariant reasons behind variations in vaccination timing (anti-vaccine movements might be stronger in one state relative to another). Year fixed effects capture year specific but state-invariant factors behind differences in vaccination timing.

Model 1 is a simple two-way fixed effects model, capturing the effect of the policy on the vaccination timing of the average population. The coefficient of interest in Model 1 is $\boldsymbol{\beta_1}$. Model 2 seeks to capture heterogeneous effects of the policy on poorer families. The coefficient of interest in Model 2 is $\boldsymbol{\beta_2}$. The vector $\mathbf{X}$ with coefficients $\boldsymbol{A}$ consists of dummy variables for the treatment, for post-treatment, for poverty, for two-way interactions, and for mother's education.

VI. Data & Descriptive Statistics

We use the publicly available National Immunization Survey (NIS) to conduct our analysis. The NIS is a series of phone surveys conducted in households with children between 19-35 months.

[17] We also perform a third analysis, using racial and marital status of the mother. Our results suggest that the policy did not have a differential effect on the outcome variables by race or mother's marital status. The coefficients of interest are either statistically insignificant or close to zero. We here omit the results of this analysis but it is available to the reader upon request.

The surveys are sponsored by the National Center for Immunizations and Respiratory Diseases (NCIRD) of the Centers for Disease Control and Prevention (CDC). The benefit of using this dataset is that it is nationally representative, has information on the family including income and mother's education, and has been conducted every year since 1994.

The treated group includes children born after the policy was implemented in California. The control group includes children born in states without PFL. One of the limitations of using the public dataset is that the exact date of birth of the child is not provided. Selected children are grouped into three age categories: 19-23 months, 24-29 months, 30-35 months. For proper treatment assignment, the child must be born after July 1st, 2004. The treated group includes children who are 19 to 23 months in 2006 and all children in 2007 onwards born in California. Since the publicly available dataset does not provide the exact date of birth of the child, assignment of treatment status becomes challenging. To address this issue, we drop the observations that are difficult to assign treatment status.

We use the vaccination and immunization module to test how access to a state mandated paid family leave affects immunization timing of infants. Looking at Table 1, we see that 90 percent of infants are late on at least one vaccine at two months of age compared to 89 percent in the control states.[18] In California, 83 percent of infants are late on at least one vaccine given before four months, compared to 82 percent of infants in the control states. It is also interesting to see that Hepatitis B-2 vaccine and Polio-2 vaccine are given to more children on time than *DTP-2* and *HIB*. In particular, the second dose of Polio vaccine is given on time more in California and control

[18] This difference is statistically significant. However, DTP-2, HIB-2 are not statistically different in terms of the average.

states than the other vaccines in Table 1.[19] This might be due to the serious nature of polio on the health of infants.

There are several positive benefits associated with early childhood vaccinations. These include a lower mortality rate of children below five, a higher attainment of human capital through lower absenteeism, and a higher productivity in adulthood because of better health and/or human capital formation (Bloom, et al. 2000; Grimwood, et al. 2000; Bloom, et al. 2003; Bloom, et al. 2004; Bloom, et al. 2005; Bloom, et al. 2007; Barnighausen, et al. 2008; Bloom, et al. 2014; Barnighausen, et al. 2011). Furthermore, the effect of vaccinations on herd immunity is well documented. The spread of communicable diseases such as Hepatitis B, whooping cough, and Haemophilus Influenzae Type b (*HIB*) can be prevented through vaccination if a sufficient number of people in a given population are vaccinated.[20] Given these positive effects, it is surprising as to why so many infants are vaccinated late.

There might be three reasons behind the late vaccination. First, parents may not believe that delaying vaccination by a few weeks is bad. In fact, they may even believe that it's good to delay immunization so that their child is not overwhelmed by potential side effects. Second, parents might face time and/or financial constraints which forces them to delay vaccination. Third, the doctors sometimes will delay immunization voluntarily if the child is on steroids or sick. Depending upon the population in a given state, we might see all three of the above factors playing a part in delaying vaccination. If the first and third factor plays a major role in delaying vaccination, then PFL will not be able to mitigate the problem of delayed vaccination. However, if time or

[19] In California, 21 percent of infants receive the 2nd dose of Polio vaccine (Polio-2) late compared to 19 percent in the control states.
[20] See "Economic Evaluation of Vaccination Programs: the Impact of Herd-Immunity" by Brisson M, Edmunds WJ (Med Decis Making. 2003 Jan-Feb;23(1):76-82) for a detailed study of how herd immunity works from an economic stand point.

financial constraints play a major role, then PFL can be used as a policy to mitigate immunization delays.

Figure 1 plots the proportion of infants being immunized late with *HepB-2* vaccine (Hepatitis B dose 2), usually given between 1 and 2 months. The figure clearly shows a striking decline in the proportion of late vaccine receivers in California, while the proportion of late vaccine receivers stays constant over time. Before 2006, the proportion of infants receiving a late vaccine is higher in California than in the other states. However, by 2010, the proportion is similar to that of the non-California average. The empirical question is whether this decline is caused by PFL mandates in California or whether something else is responsible for this decline.

VII. Results

A. MAIN RESULTS

Table 2 presents results of Model 1 for late vaccination of infants for three particular vaccines: *HepB-2*, *DTP*, and *HIB*. Column 1 of Table 2 suggests that the policy decreased the likelihood of vaccinating an infant late with *HepB-2* by 5 pp relative to that in the control states due to the policy. Columns 2 and 3 suggest that the likelihood of vaccinating an infant late in California with DTP or HIB shots is reduced by approximately 1.4 pp because of the policy.

Table 3 presents results of Model 2. The results suggest a stronger effect of the policy on poorer families relative to families above the poverty line and others in the control state. In particular, the probability of late vaccination of an infant from a poor family in California is reduced by 5 pp to 7 pp relative to infants born to other families in California, and similar infants from other states that lack paid parental leave policy.

We conduct a series of placebo tests to assess the validity of our results and the parallel trend assumption. We first conduct a soft placebo test in which we drop California from our analysis and select states that are similar to California in terms of size, population, demography, political leaning. The idea of this test is to verify whether results are similar for states that resemble California but did not pass PFL legislation. The plausibility of our results becomes weak if the estimates using these placebo "California" states are similar to that of the real California. Specifically, violation would occur if the coefficients using the placebo states are statistically significant and have similar or higher magnitudes compared to the true estimate. In an ideal state of the world, the placebo states should display a coefficient of zero since the policy was not implemented there.

Given how unique California is, finding suitable states resembling the state can be particularly challenging. We pick New York, Washington, Texas, Florida and Massachusetts as our placebo states. New York and Texas are as densely populated as California. They also exhibit similar diversity in terms of immigrant population, ethnicity and race[21]. New York, Washington and Massachusetts also had been contemplating implementing PFL legislation similar to that of California. Indeed New York passed PFL legislation in April, 2016, which went into effect beginning January 1st, 2018. Moreover, these states were also used as placebo states by other researchers such as Rossin-Slater et al. (2013), and Das & Polachek (2015) in their papers analyzing the effect of Paid Parental Leave on labor market outcomes.

The results are presented in Tables 4 and 5. None of the coefficients of interest are statistically significant. For HepB-2, the coefficient of ***Treat*Post*** from Model 1 is positive when the true

[21] It should be noted however, that California has a lower proportion of black individuals relative to that of NY and TX.

estimate is negative. The coefficients in Table 4 are either very close to zero or have the incorrect sign. The coefficients for ***Treat*Post*Poor*** are all statistically insignificant and the point estimates are close to 0.

Some might question whether using New York, Washington, Texas, Florida and Massachusetts are good placeboes for California. To address this criticism, we conduct a strong placebo test. Here we drop California from our sample, and pretend that the remaining control states are placebo California, one by one. For instance, we pretend that Alabama is a placebo for California, then we repeat the analysis by pretending Alaska is a placebo for California, then Arizona and so on. We plot a distribution of the placebo coefficients and compare them with the true coefficient for California. Our results will be weaker if the true estimate lies in the 90 percent confidence interval of the distribution.

Figures 2 to 5 illustrate the kernel distribution of placebo coefficients and compares the true coefficients of interest from Tables 2 and 3 to the placebo distribution. All four figures show that the true coefficient is outside the 95 percent of the kernel distribution, suggesting that our results are quite plausible. To be precise, Figure 2 shows that only 3 states have magnitudes greater than the true coefficient (-0.051). Figure 3 shows that only 2 states have magnitudes greater than the true coefficient (-0.0683). Figure 4 shows that only 1 state has a placebo coefficient greater in magnitude than the true coefficient (-0.06). Finally, Figure 5 shows that only 3 states violate the placebo test. Therefore, performing the strong placebo test suggests that the policy indeed had some effect on vaccination timing, and that the results are not driven by chance or specification error. We do not show the placebo test distributions for ***Treat*Post*** in Model 2 for *DTP* and *HIB* because the point estimates are close to zero and are statistically significant only at the 10 percent level.

We perform another robustness check to determine whether the results are being driven by the policy. Since the California PFL policy provides only six weeks of leave with partial wage replacement, we shouldn't see any effect of this policy for vaccines that are given later in an infant's life. Women, who have access to temporary leave under other existing policies such as Temporary Disability Insurance (TDI) or State Disability Insurance (SDI), usually take leave through California's policy after they have already exhausted their usual leave period. On the other hand, men take leave around the time of birth of their child. For long term vaccines, we use HepB-3 and Polio-3, which are given after the child is six months old. Table 6 presents the results of this analysis. The table shows that the policy had no such effect on the timing of getting these vaccines. In other words, all the coefficients of interest have a point estimate close to zero.

B. SYNTHETIC CONTROL METHOD & LOGISTIC REGRESSION

We repeat our analysis using the synthetic Control Method (SCM) as developed by Abadie, et al. (2010). The benefit of using SCM over a simple difference-in-difference estimation technique is that the pool that resembles the treated group in terms of predictor variables gets assigned a higher weight. Therefore, states that are more similar to California are weighed more heavily. Since the synth command requires that all units be balanced, we drop 2006 since some selected children in California are treated while others are not.

Figure 6 illustrates that there is a decline in the proportion of infants being vaccinated late with *HepB* in both the synthetic control group as well as the treated group (infants born in California). However, the decline in the treated group is larger than that in the synthetic control group. This additional decrease is because of the policy. Similarly, figures 7 and 8 illustrate that over time, the proportion of infants receiving the vaccines *DTP* and *HIB* late is increasing while that for the treated group stays around the pre-policy trend. These results strengthen our original conclusion that the policy indeed reduced the likelihood of infants receiving vaccine late.

Finally, we run a logit regression model in order to compare our results to the linear probability model. We present both the odds ratios and marginal probabilities in Tables 7 and 8. The results similarly indicate that the policy reduced the likelihood of delayed immunization. In particular, the policy reduced the likelihood of an infant from a poor household receiving DTP and HIB late by almost half.

VIII. Conclusion

The issue of paid family leave is becoming relevant in political discussions. It is imperative to study how paid family leave affects economic outcomes and household decisions regarding child development. This paper uses a unique dataset to test how short duration paid leave, mostly taken for purposes of bonding with a newborn baby, affects vaccination timing. We test our main hypothesis that parents may be delaying vaccinations because of time constraints. We use a unique dataset obtained from the Centers for Disease Control and Prevention. Employing a difference-in-difference estimation technique, our results suggest that short duration paid parental leave policies (like the one in California) can improve on time vaccination of infants. Our results also suggest that the policy had a greater impact on poorer families. This finding suggests that poor families might have less access to other facilities such as employer provided leaves, family support, or the ability to afford formal care that help them with child rearing responsibilities.

There are various reasons to think that the estimates in our paper are a lower bound of the true effect of paid parental leave policies on vaccination timing. Firstly, parents may not take the full leave time because the wage replacement is not that high. Since we do not observe the actual leave taken by the parents, our identification relies on the assumption that everyone having children after the policy implementation in 2004 is eligible to take leave for six weeks and to receive partial wage replacement. Secondly, some people that are eligible for leave may not be aware of the policy (Ruth, 2008). Finally, California's PFL policy does not offer job protection. While some type of job protection is offered in California through CFRA (California Family Rights Act) and FMLA (Federal Family and Medical Leave Act), the six weeks additional leave under PFL only provides partial wage replacement. This might lead some parents not to take any leave or take fewer days off under the policy.

There are a few interventions that can potentially fix these. Companies can explicitly inform their employees about PFL during training to reduce lack of awareness. Alternatively, independent organizations can advertise the PFL provisions. However, both these options involve some cost. Also, legislators can increase the benefit cap (which might eventually lead to raising payroll taxes) or make PFL include job protection. Amending the current policy to cover job protection might be the least expensive solution compared to the alternatives listed above.

Importance from a policy standpoint is to investigate which factors contribute to the inequality in health outcomes among children from different socio-economic backgrounds. Our results suggest that some parents delay immunization due to time constraints. These time constraints might be due to work related obligations. Furthermore, the problem can be exacerbated if the parents are working in industries that do not have provisions for parental leave. Our paper finds that a policy providing six weeks leave with partial wage replacement has a significant effect on improving immunization of infants on time, with a stronger effect on poorer families. Baseline estimates suggest that 78.3 percent, 92.2 percent, and 91.5 percent of poorer households report delaying vaccinating the infants with *HepB*, *HIB*, and *DTP* vaccines respectively. Compared to the baseline estimates, our results suggest that the policy reduced the likelihood of delayed vaccination for infants born to poor households by 9 percent, 5.4 percent, and 6.6 percent below the baseline for *HepB*, *HIB*, and *DTP* vaccines respectively.

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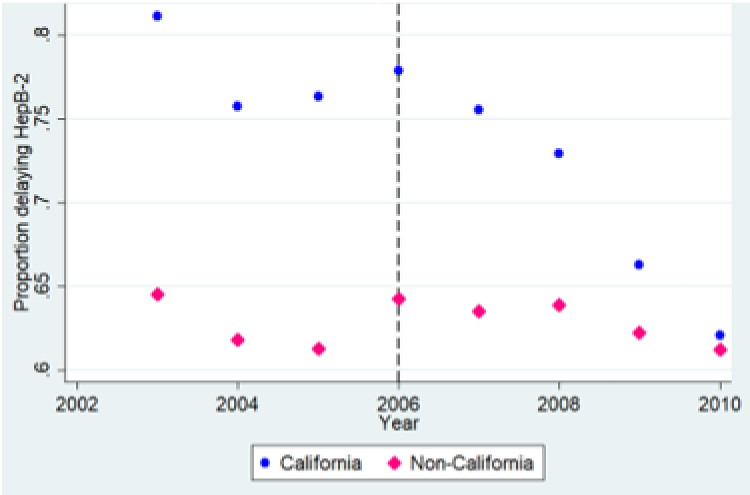


Figure 1

Note: Figure 2.1 plots the average proportion of children in who are vaccinated late with Hepb-2. We see that for the control group (non-California), the trend is quite stable across time. However, for the treated group (California), the proportion who receive vaccination late drops 78 percent in 2006 to 62 percent in 2010. Moreover, in 2010, the proportion of infants vaccinated late with Hepb-2 in California and non-California states is not that different.

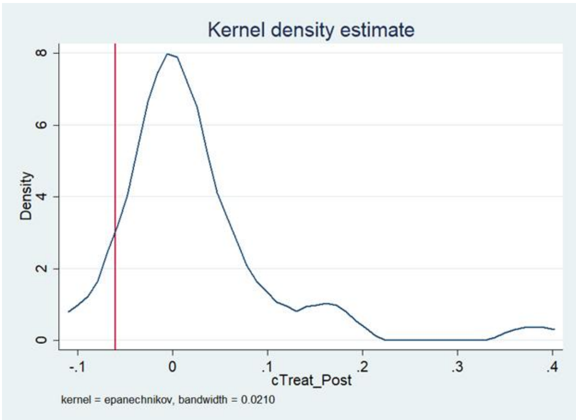


Figure 2

Note: Kernel distribution of ***Treat*Post*** (Model 1) using all 48 states as placebo California where late vaccination of HepB-2 is the outcome variable, using Model 1. Only 3 states violate this placebo test.

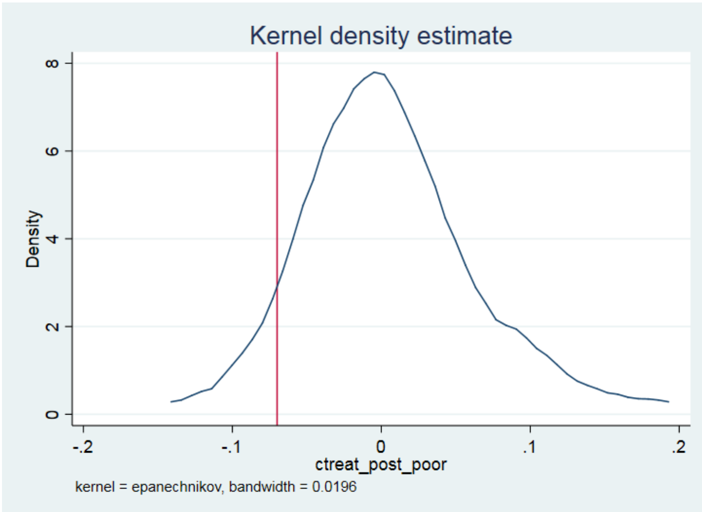


Figure 3

Note: Kernel distribution of ***Treat*Post*Poor*** using all 48 states as placebo California where late vaccination of HepB-2 is the outcome variable. 2 states violate this placebo test.

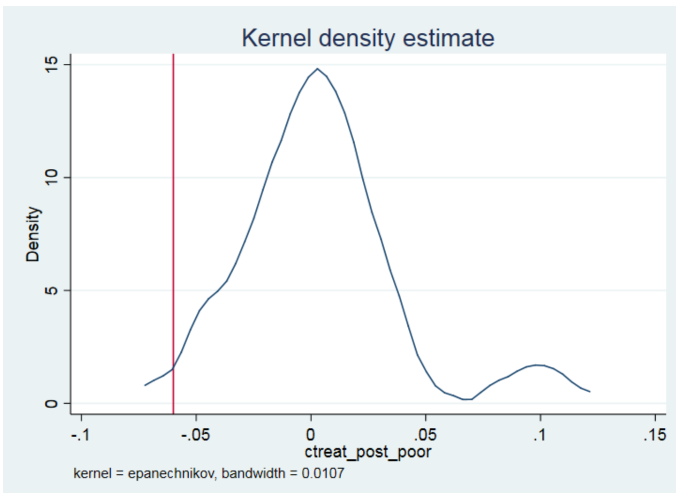


Figure 4

Note: Kernel distribution of ***Treat*Post*Poor*** using all 48 states as placebo California where late vaccination of DTP is the outcome variable. 1 state violates this placebo test.

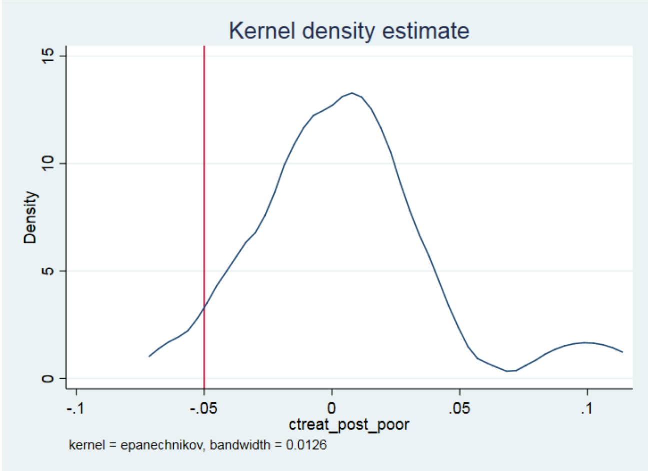


Figure 5

Note: Kernel distribution of ***Treat*Post*Poor*** using all 48 states as placebo California where late vaccination of HIB is the outcome variable. 3 states violate this placebo test.

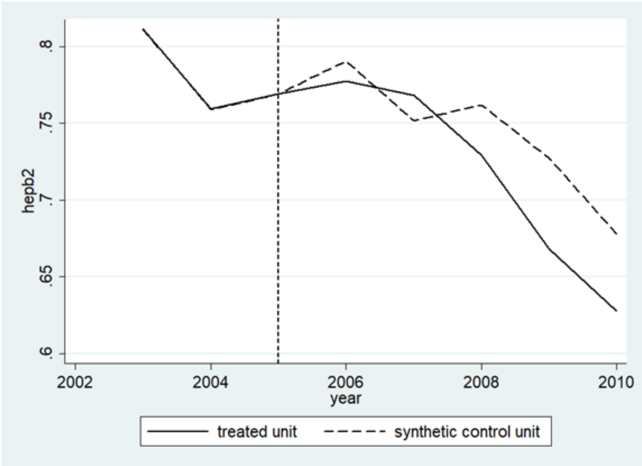
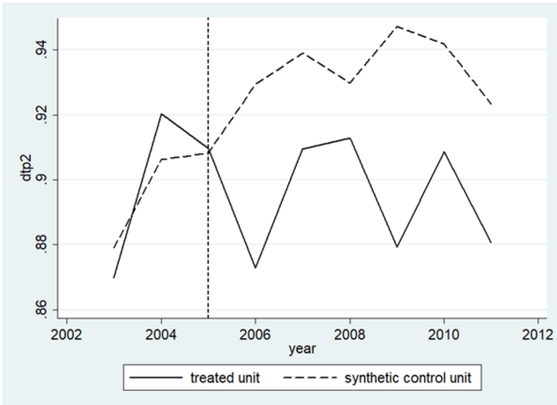


Figure 6

Note: The results suggest a 15-pp decline from 2006 to 2010. The synthetic control group declines by 11 pp. This suggests that the policy had an additional 4-pp negative effect.



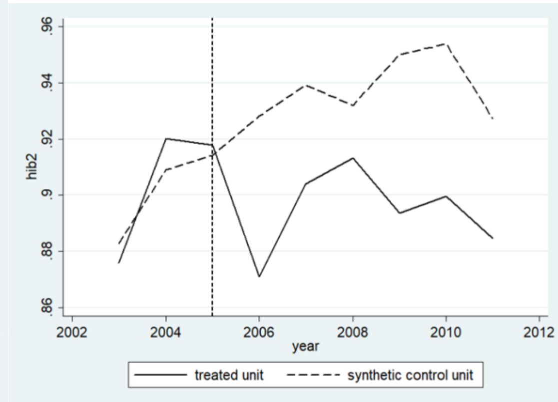


Figure 7 (Left Panel): DTP Vaccine

Figure 8 (Right Panel): HIB Vaccine

Table 1: Descriptive Statistics

	California		Non-California	
Mother's Education Level:				
Single	0.23	(0.0033)	0.21	(0.0008)
No High School	0.20	(0.0031)	0.12	(0.0006)
High School graduate	0.23	(0.0033)	0.24	(0.0008)
Non-College graduate	0.19	(0.0031)	0.22	(0.0008)
College graduate	0.38	(0.0038)	0.42	(0.0009)
Poor	0.24	(0.0033)	0.18	(0.0007)
Vaccination Outcomes:				
DTP	0.87	(0.0033)	0.88	(0.0007)
Hib	0.87	(0.0033)	0.88	(0.0007)
HepB	0.69	(0.0045)	0.61	(0.0001)
Late on at least 1 vaccine given by 2 months	0.89	(0.0031)	0.88	(0.0007)
Late on at least 1 vaccine given by 4 months	0.79	(0.004)	0.79	(0.0009)

Note: The standard error of mean is recorded for each variable in parenthesis, next to the mean. The vaccination outcomes reflect the mean of children in California vs non-California states who are not vaccinated on time. Source: CDC, NCRID and NCHS, National Immunization Survey. Data is merged from 2000-2011.

Table 2: Model 1 Results for Vaccines

VARIABLES	(1) HepB-2	(2) DTP	(3) HIB
Treat*Post	-0.0510***	-0.0138***	-0.0136***
	(0.00817)	(0.00213)	(0.00224)
Observations	174,242	175,401	174,954
StateFE	Yes	Yes	Yes
YearFE	Yes	Yes	Yes

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1
Source: CDC NCRID and NCHS (2000-2011)

Table 3: Model 2 Results for Vaccines

VARIABLES	(1) HepB-2	(2) DTP	(3) HIB
Treat*Post	-0.0239**	0.00935***	0.00831***
	(0.00951)	(0.00276)	(0.00244)
Treat*Post*Poor	**-0.0683*** **	**-0.0599*** **	**-0.0479*** **
	(0.0159)	(0.00464)	(0.00442)
Observations	159,431	160,507	160,121
StateFE	Yes	Yes	Yes
YearFE	Yes	Yes	Yes

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1
Source: CDC, NCRID and NCHS (2000-2011)

Table 4: Placebo for Model 1 Results for Vaccines

VARIABLES	(1) HepB-2	(2) DTP	(3) HIB
Treat*Post	0.0502	0.00560	0.00783
	(0.0425)	(0.00846)	(0.00775)
Observations	174,242	175,401	174,954
StateFE	Yes	Yes	Yes
YearFE	Yes	Yes	Yes

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1
Placebo states used are Massachusetts, New York, Texas, Florida, Washington.
Source: CDC NCRID and NCHS (2000-2011)

Table 5: Placebo for Model 2 Results for Vaccines

VARIABLES	(1) HepB-2	(2) DTP	(3) HIB
Treat*Post	0.0521	0.00427	0.00575
	(0.0447)	(0.00845)	(0.00836)
Treat*Post*Poor	-0.00882	0.00194	0.000463
	(0.0233)	(0.00500)	(0.00406)
Observations	159,431	160,507	160,121
StateFE	Yes	Yes	Yes
YearFE	Yes	Yes	Yes

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1
Placebo states used are Massachusetts, New York, Texas, Florida, Washington.
Source: CDC, NCRID and NCHS (2000-2011)

Table 6: Long-term Vaccines

VARIABLES	(1) HepB_3	(2) Hepb_3	(3) Polio_3	(4) Polio_3
Treat*Post	0.00028	0.00678***	0.00928***	0.00501***
	(0.00178)	(0.00172)	(0.00288)	(0.00177)
Treat*Post*Poor		-0.00090***		-0.00275
		(0.00293)		(0.00304)
Observations	220,029	108,309	227,728	109,220
R-squared	0.005	0.006	0.013	0.013
StateFE	Yes	Yes	Yes	Yes
YearFE	Yes	Yes	Yes	Yes

Robust standard errors in parentheses, clustered at the state level.
*** p<0.01, ** p<0.05, * p<0.1
Note: We used other long-term vaccines and found no strong significance of the policy. The coefficients are very close to zero (barely 1 pp effect size). Source: CDC, NCRID and NCHS, National Immunization Survey. Data is merged from 2000-2011.

Table 7: Logit Regression Results for Model 1 Results (Odds-Ratio Reported)

VARIABLES	(1) HepB-2	(2) DTP	(3) HIB
Treat*Post	**0.762***** **(0.0280)** [-.0531]	**0.871***** **(0.0191)** [-.0134]	**0.870***** **(0.0223)** [-.0137]
Observations	174,242	175,401	174,954
StateFE	Yes	Yes	Yes
YearFE	Yes	Yes	Yes

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1
Source: CDC, NCRID and NCHS, National Immunization Survey. Data is merged from 2000-2011. The marginal effects are given in [] brackets below the standard errors.

Table 8: Logit Regression Results for Model 2 (Odds-Ratio Reported)

VARIABLES	(1) HepB-2	(2) DTP	(3) HIB
Treat*Post	0.870*** (0.0375) [-0.028]	1.104*** (0.0299) [0.0095]	1.082*** (0.0276) [0.008]
Treat*Post*Poor	**0.715***** **(0.0512)** [-0.066]	**0.467***** **(0.0223)** [-0.067]	**0.545***** **(0.0241)** [-0.052]
Observations	159,431	160,507	160,121
StateFE	Yes	Yes	Yes
YearFE	Yes	Yes	Yes

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1
Source: CDC, NCRID and NCHS, National Immunization Survey. Data is merged from 2000-2011. The marginal effects are given in [] brackets below the standard errors.