STOPPING THE BIOLOGICAL CLOCK: INFERTILITY TREATMENTS AND THE CAREER-FAMILY TRADEOFF

Abstract

Career women who want children weigh two competing forces. The fact that fecundity declines with age pushes them to have children earlier. On the other hand, career motives encourage them to delay childbirth. This paper examines the impact of recently developed infertility treatments, which relax the biological time constraint on childbearing. I use state time-series/cross-section variation in insurance coverage for infertility treatments to show that improved access to the technologies does result in changes in fertility outcomes. Most notably, I find that improved access to infertility treatments increases the incidence of multiple births and the birth rate among older women. I then use the same technique to explore the possibility that an extended biological clock will lead to changes in women's investment in human capital and pursuit of careers. Mandating insurance coverage increases labor force participation for women under 35, and decreases participation for women over 35 while increasing their wages. Therefore, it appears that extending the biological window for childbearing does improve the career-family tradeoff for women by making delay less costly.

J130 - Fertility; Family Planning; Child Care; Children; Youth

J200 - Time Allocation; Work Behavior; Employment Determination and Creation; Human Capital; Retirement:

I180 - Health: Government Policy; Regulation; Public Health

I. Introduction

Since women began entering the U.S. labor force in large numbers in the second half of the twentieth century, many working women have struggled to balance their careers with their desire to bear and raise children. Having a child almost always means career interruptions at the time of the birth, and raising a child requires time, money, and other resources that otherwise might have been used to invest in human capital and increase labor productivity. An important component of the career/family tradeoff is the timing of births, which for career women is largely a struggle between two competing forces. On the one hand, decreasing fecundity by age pushes women to have children earlier. On the other, career motives encourage women to have children later. This paper will explore the possibility that recently developed infertility treatments and technologies, which relax the biological time constraint on childbearing, might allow more women to increase their investments in human capital and pursue careers.

Previous research shows that delaying childbirth until later in a career leads to greater lifetime earnings. Ellwood, Wilde, and Batchelder [2004] show the effects of the timing of childbirth on wages for women of varying skill levels. Women in the highest skill group appear to be on similar wage trajectories until the birth of their first child. Remarkably, wage profiles in all cases flatten out at approximately the point of the birth. They also find that the pay penalty increases with time since the birth, resulting in a 15 to 20 percent decrease in pay after 10 years. Finding no evidence of decreases in women's wages prior to the first birth, they conclude that the drop in wages can not be explained by women deciding to have a child when their wage growth was slowing.

However, delaying childbirth can be risky. One of the most binding constraints women face is the biological time constraint on bearing children, or the "biological clock." Menken,

Trussell, and Larsen [1986] find that, among women not using contraception, the percent of couples who remain childless varies by age at marriage: 6% for couples married at age 20-24, 9% at ages 25-29, 15% at ages 30-34, 30% at ages 35-39, and 64% at ages 40-44. As the probabilities of becoming pregnant and carrying a pregnancy to term fall, so does the probability of having a healthy child. Van Noord-Zaadstra, et al [1991] claim that the probability that a non-contracepting woman conceives and delivers a healthy baby falls by about half from age 25 to age 35.

Goldin [2004] studies groups of women from the twentieth century by cohort, and analyzes their career and family decisions. She finds that the cohort graduating from college between 1966 and 1979 sought to have both a career and a family, either by having a career first or by trying to have both at the same time, but only 13 to 18 percent of these women were successful in doing so[1]. However, the most recent cohort in Goldin's analysis, which graduated from college between 1980 and 1990, fares significantly better than previous cohorts with between 21 and 27 percent achieving both a career and a family. Could it be that an extended biological clock has helped more women from this cohort than from previous cohorts reach the goal of "having it all?" And could future generations see further progress as fertility technologies become more widely available? [2]

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1. Goldin defines a career as "reaching an annual income (or hourly wage) level greater than that achieved by a comparable college graduate man who was at the 25th percentile of the male annual income (or hourly wage) distribution" for two to three consecutive years in the woman's late thirties or early forties.

2. Goldin and Katz's [2002] research on the birth control pill provides another notable example of the effect of technological improvements in fertility control on women's economic outcomes.

In fact, the possibility of using infertility treatments to delay childbirth seems to be especially appealing to career women, for reasons I will discuss in this paper. In a survey of young, high-achieving women, 89% believed that infertility treatments would allow them to get pregnant into their forties [Hewlett 2002].[3] If women are planning on using infertility treatments to delay children, then we would expect to see changes in labor force participation profiles, with interruptions occurring later in their careers. The effects of delaying on wages are not as clear. I consider the literature on the timing of career interruptions, and discuss the circumstances under which delaying could increase wages for working women.

Empirically, I use U.S. state-level variation in laws which mandate insurance coverage for infertility treatments to analyze both changes in birth patterns and changes in women's career patterns. I implement a time-series/cross-section estimation strategy with state-level fixed effects and a flexible time trend, using birth certificate data from the CDC's natality detail files from 1982-1999, as well as labor force and demographic information from the March Current Population Survey for the same years.

I find that in states where infertility treatments are more available due to insurance coverage, there is an increase of about 23% in the rate of multiple births, which suggests increased use of infertility treatments. There is also evidence of an increase in the birth rate for women ages 35-44 in states with mandated insurance coverage. At the same time, the percent of young women who are mothers in these states decreases, indicating that at least some of the

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3. There is some evidence that women's perceptions of the effectiveness and availability of these treatments are overly optimistic [Hewlett 2002]. However, in order to see the career effects discussed in this paper, it is only necessary that women believe that infertility treatments will allow them to delay—whether or not that is actually the case.

increase in the birth rate for older mothers is due to women choosing to delay childbirth.

Having shown that fertility decisions are affected by the availability of infertility treatments, do we observe changes in career patterns like the ones discussed above? I show a shift in the age-labor force participation profiles for women in states with a mandate. Women in these states are relatively more likely to participate in the labor force before age 35, but less likely after this point. I also find some evidence of increased wages for college-educated working women in their late 30s and early 40s among women covered by a mandate.

This paper is organized as follows. In the next section, I discuss the theory on the effects of interruptions and their timing on careers. The third section provides a brief background on infertility treatments, and the fourth presents the estimation strategy. The fifth presents empirical results that show the effect that improved access to these therapies has on family and labor force outcomes. The last section concludes.

II. The Theory of Career Interruptions

Before looking for empirical evidence that infertility treatments improve women's career outcomes by allowing them to delay childbirth, it is necessary to examine the theory underlying this hypothesis. Relaxing the biological time constraint on childbearing expands the choice set and therefore should be welfare-improving in any economic model of an individual's career and family decisions, because the option not to delay is always available. However, in order to better understand the magnitude of the impact that infertility treatments might have, we would like to know the circumstances under which women might find that delaying interruptions results in higher career attainment.

A. The Case for Early Childbearing

In a Ben-Porath style model in which agents choose between investment and paid work,

individuals choose to invest early partly because the opportunity cost of time spent out of the labor force (wages) is lowest when human capital is also lowest. This model could be generalized to allow women a third option—time at home to have a child. If women are allowed to continue to invest when they have a child so that human capital and thus earnings increase over time, the opportunity cost of foregone wages is lowest early in the career. This structure would motivate early childbearing.[4]

Consider also the case in which wage growth does not depend on work experience—for example, if there is secular wage growth in the economy, as in Miller [2005], or if there are returns to age that are independent of work experience.[5] If wages increased faster than discounting, women would maximize lifetime earnings by interrupting their careers early when wages are low. Empirical evidence that it may be better to have children early can be found in Hotz, McElroy, and Sanders [1996] and Hotz, Mullin, and Sanders [1997]—though these studies are more relevant for lower-skilled women. Because investment is less important in low-skilled careers, wage profiles for these women may be similar to the secular wage growth or returns to age models.

B. The Case for Delayed Childbearing

Empirically, the majority of the evidence suggests that higher-skilled women and women on career paths are more likely to delay children, though the direction of causality is unclear. Heckman and Walker [1990] show that for women in Sweden, both higher wages and selection of a white collar job increase the time to first birth. In Blackburn, Bloom, and

4. Results are based on a fully developed model available in an earlier version of this paper.

5. Lazear [1976] finds positive returns to age even after controlling for experience and education.

Neumark [1993], women who have children later have higher early wages than women who have children earlier, which the authors attribute to greater human capital investment on the part of delayers. Taniguchi [1999] uses NLSY data and finds that independent of work experience, delayed career interruptions are associated with higher wages—a result which she attributes to missed "career building" opportunities for early childbearers.

These ideas are also represented in structural economic models in which career incentives provide motivation for delaying children. In these models, at least some part of the fertility timing decision depends upon the opportunity cost of an interruption and the effect of the interruption on the lifetime earnings profile. Caucutt, Guner, and Knowles [2002] construct a model with both marriage and labor markets, in which productivity in the labor market depends on past labor supply decisions. In their calibrated model, they find that women are more likely to work early in their careers and are more likely to delay children when returns to experience are positive than when they are zero. Erosa, Fuster, and Restuccia [2002] develop a structural model of women's career outcomes that includes a fertility timing decision. Age at birth in their model "is likely to be related with tenure on the job and, thus, with the specific human capital that the female would lose if quitting the job." Because "job separation today increases the likelihood of a job separation in the subsequent job," early career interruptions for women are costly and the effects are long lasting.

Also, to return to the Ben-Porath model, the above result that women should have children early was conditioned on their ability to continue to invest during career interruptions. In reality, this may not be feasible, possibly because having a family competes with resources necessary for investment, such as purchased inputs or time. Also, credit constraints could mean that women could not afford to have a child if they were not working. On the labor market side,

investment and work may be complements, so that it is not feasible to invest with little or no work. Learning-by-doing, for example, would require that women do a minimum amount of work to be able to invest.

In such a model, women who interrupt their careers early will have less human capital after their interruption than women who have not yet had an interruption. Therefore, even though they may have the same human capital in the end, lifetime earnings are higher for delayers. It is the foregone investment early in the career that is most costly in this case.

Finally, notice that some basic assumptions within any economic model can motivate delayed childbearing. For example, discounting of future income in a model makes women more likely to delay children, since women would rather be out of the workforce when earnings are more heavily discounted. Miller [2005] finds that a constant depreciation of human capital during an interruption (independent of the timing of the interruption) could result in delayed childbearing, since women who delay will work longer without any depreciation. This idea is similar to the above story in which women forego investment during an interruption. She also shows that if women experience decreased returns to experience after a birth, perhaps because of "reduced opportunities for training and promotion," women will have flatter wage profiles after a birth and will therefore have higher lifetime earnings if they delay children.

I now turn to a brief discussion of fertility technologies and the role that they might play in allowing women to delay.

III. Infertility treatments

A. Development and Use

Between 1990 and 1997, pregnancy rates declined for women in their twenties, while rates for women ages 30-34 increased 3 percent and those for women ages 35-39 increased by 9

percent [Ventura et al 2001]. For women ages 40-44, the rate of pregnancy increased by 21 percent. Over the same period, recently developed technologies have been utilized that help relax the biological time constraint on bearing children. What role have these technologies played in the trend toward delayed childbearing?

Infertility treatments are medical technologies that help women and couples with infertility problems conceive a child. Services range from counseling sessions and tests to fertility drugs and invasive surgical procedures. Surgeries include assisted reproductive technologies (ARTs) such as in vitro fertilization (IVF), in which an egg is taken from the woman's ovaries, fertilized, and then placed in her uterus. IVF was first successfully performed in England in 1978, with the first successful procedure in the U.S. in 1981.

Infertility treatments have become increasingly effective over the past twenty years. Since 1995, the Center for Disease Control has required clinics that perform ARTs to report success rates for the various procedures [CDC 2003, CDC 2005]. By 2003, the data shows that for cycles using fresh embryos from nondonor eggs, the procedure had almost a 45% success rate for women under 35, where success is defined as a live birth.[6] For women in their late thirties and early forties, success rates were approximately 25% and 10%, respectively.

The rate of use of infertility treatments has also increased. The CDC reported that in 2003, 122,872 ART cycles were performed in 399 clinics in the U.S., resulting in 35,785 live births and 48,756 infants. Figure I shows the increase in the use of ARTs, as well as the outcomes, between 1996 and 2003.

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[6] Note that 20% of fertile couples will experience a live birth from a natural cycle, as reported by the American Society of Reproductive Medicine [2006]. Also, 75% of AFT procedures use fresh embryos from nondonor eggs.

Finally, the 1995 National Survey of Family Growth provides insight into the type of woman who uses infertility treatments. 15.4% of all women and 22.9% of women ages 35-44 report having received infertility services. Of those, the majority received advice, tests, or fertility drugs. Only 1% of women nation-wide received ARTs, though for women 35-44 with no children this figure is 5.3%. Older women are much more likely to have sought all types of infertility treatments, particularly those who are married, those who have a college degree, and those with high incomes. Among women with only a high school degree, 1.1% have used ARTs, but this figure doubles to 2.2% for women with a bachelor's degree. This suggests that it is in fact career women who are making use of infertility treatments. Certainly infertility treatments are more available to these women, who are more likely to be able to afford these services and are also more likely to be near clinics (as clinics are more concentrated in urban areas). But these figures also support a story in which career women have stronger incentives to try to extend their biological window for having children.

B. Infertility treatments and Fertility Outcomes

Because the widespread use of infertility treatments is a relatively recent phenomenon, little work has been done on the fertility effects of their development. A few studies have examined the impact of infertility insurance mandates on fertility outcomes, as I do below. Hamilton and McManus [2005] analyze the effect of the insurance mandates on the market for infertility treatments. They show that insurance mandates increase the number of cycles performed and decrease the number of embryos transferred per cycle. Similarly, Schmidt [2005] finds that the presence of a mandate increases birth rates for women over 35 by about 32%. Bitler [2005] examines the effect of the mandates on infant health, and finds that insurance mandates are associated with an increase in twin births, as well as slightly less healthy twins as

measured by gestation, weight, and 5-minute APGAR scores.

Evidence of increased use of infertility treatments can be obtained using data from the National Center for Health Statistics' Natality Detail Files from 1982 to 1999. The data provides records for all births in 51 U.S. states for every year (including the District of Columbia), with the exception of a few states that report 50% of births in 1982 and 1983. Each record contains information such as the place of the birth, mother's state of birth and state of residence, mother's age and race, and in most cases mother's and father's education levels.

The data set does not allow us to observe infertility treatment use directly. However, we do observe the number of children per birth—that is, we know if the birth resulted in twins, triplets, or a higher-order multiple birth. Multiple births are strongly correlated with infertility treatment use. In fact, the CDC reports that 35.4% of live births resulting from ART produced two or more infants, compared to 3% for the general population [CDC 2003]. Furthermore, they find that over 40% of triplet and higher-order births in 1997 were a result of ARTs, and another 40% were due to fertility drugs [CDC 2000].

Figure II shows the number of infants born per first birth in the U.S. between 1982 and 1999 (so that a singleton first birth is recorded as 1, twins as a first birth as 2, etc.).[7] We might expect multiple births to differ by age for two reasons. One, older women are naturally more likely to have multiple births. Two, older women are also more likely to experience infertility and therefore seek infertility treatments. Note that in the early 1980s, before the widespread use of infertility treatments, rates of multiple births were relatively low, with older women slightly more likely to experience multiples. However, by 1999, the number of additional children per

[7] Only white women are included because the likelihood of multiple births varies by race. Limiting the sample to white women avoids selection effects.

first birth is over 3 times the 1982 level for women over 35. Despite the increases for women in older age groups, the number of children per first birth for women under 25 has remained almost constant. These results suggest that older women are turning to infertility treatments to allow them to have a child later in life.

The number of multiple births also varies by education level. This strongly suggests infertility treatment use—ceteris paribus, there is no reason that multiple births would vary by education. The results are presented in Figure III. Because older women and women from certain race groups are more likely to have multiple births, and because the age and racial compositions of the groups may have changed over the period, I confine the sample to white women whose first birth occurs after age 30. Note that in the early 1980s, before widespread use of infertility treatments, the three groups were similar in the number of children per first birth. By 1999, the number of infants per first birth has grown substantially for women with a high school diploma or some college and even more for women with at least a college degree. This suggests that infertility treatment use is increasing in education, confirming the findings of the 1995 NSFG.

These increases in births to older women and increases in multiple births indicate that infertility treatments have affected women's family decisions in the United States, particularly for more educated women. I now look to confirm these results empirically.

IV. Estimation

A. Identification Strategy

To estimate the effects of fertility technologies on women's birth and career choices, we would like to be able to compare women with access to the technologies to those without such access. One possibility would be to look at women in years before the technologies were

developed and compare them to women after they became widely available. However, as described above, the journey from invention to widespread use took many years and happened gradually. Therefore, the women in the two groups would differ significantly in many ways that would certainly not be due to their access to fertility technologies alone. Alternatively, one could look at geographic access to clinics. Particularly in the years immediately after the technologies were developed, clinics performing infertility treatments were concentrated in urban centers. Unfortunately, we again have the problem of important differences in the treatment and control groups—women in large cities likely have unobserved characteristics that make them different from women in less-urban areas, and those characteristics could be correlated with their family and career choices. Furthermore, women who are willing to pay tens of thousands of dollars for fertility technologies are likely willing to pay the additional but relatively small cost of a trip to a clinic.

This last point, however, brings us to a promising source of variation in FT access—cost. The monetary costs of conceiving using fertility technologies can be prohibitively high for many women. Fertility treatments are generally conducted as monthly cycles, with several cycles often being necessary. In 1996, a typical cycle of ovulation-inducing drugs ranged between $200 and $3,000 [Davis 1996]. ART is much more costly at approximately $12,000 per cycle [Hewlett 2002]. In 1993, the cost per live delivery from an IVF cycle in Massachusetts was estimated to be almost $70,000, and over $100,000 for women age forty or over [Griffin and Panak 1998].[8]

Insurance plans typically cover small portions of these costs or none at all. In their 1995 national survey of employer-sponsored health plans, Foster Higgins reports that about one-fourth

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8. Cost per live delivery is the total cost of all IVF procedures done in the sample, divided by the number of live deliveries resulting from those procedures [Griffin and Panak, 1998].

of all firms nationwide with ten or more employees provide some form of infertility insurance, though the bulk of the coverage is for professional services and drug therapies and not ARTs. A study by the Alan Guttmacher Institute [1994] suggests that coverage for in vitro fertilization is not common—ranging between 14 and 17% depending on the type of coverage. Many policies explicitly exclude IVF and other fertility services. Individual policies covering infertility insurance are not available due to the obvious adverse selection issue.

However, between 1982 and 1999, thirteen states passed legislation requiring at least some coverage of infertility treatments (see Table I).[9] I classify six of these states—Maryland, Arkansas, Hawaii, Massachusetts, Rhode Island, and Illinois—as having strong "mandate-to-cover" laws, which require all non-self-insured employers to provide health plans with coverage for treatment of infertility. To be classified as having a strong mandate-to-cover law, states must have at least 35% of women covered by the law, and in vitro fertilization must be covered. In comparison, about 25% of employer-sponsored insurance plans cover any treatment of infertility, and many of those exclude IVF and other services [Nelson 1999]. Montana, Ohio, and West Virginia have laws that are described as "mandate-to-cover," but only HMOs are required to comply. In 2004, these three states exhibited HMO penetration rates of only 8.8%, 15.5%, and 10% respectively [Kaiser Family Foundation 2004]. Furthermore, the Montana and Ohio mandates exclude IVF. In New York, the law merely states that insurance companies can not deny treatment of a covered condition only because that condition results in infertility. When applied, this law was quite weak, and was strengthened in 2002, which is beyond the period covered in this study. The remaining three states only require that a policy including infertility

9. Insurance law information was obtained from Resolve [2004] and the National Conference of State Legislatures [2004].

treatment coverage be offered to employers. However, the employer may still choose a plan with no infertility insurance provision. Thus, in practice, a mandate-to-offer law still leaves the choice of infertility coverage up to the employer—much like the case in states with no law at all.

Therefore, despite the fact that not all women are covered in mandate-to-cover states, the high cost of fertility technologies has been reduced by thousands of dollars to effectively zero for many women. As a result of the decrease in costs, the average woman in these states has substantially greater access to fertility technologies than women in states with no law. Perhaps more importantly for any career outcomes, women in these states may be more aware of the availability of infertility treatments if more women around them are using them. Therefore, I will compare women in strong mandate-to-cover states to states with no such law. For the bulk of the analysis, I rely on estimates which omit the weak mandate-to-cover and mandate-to-offer states, to avoid contamination of the control group. For comparison, the tables include results that do have these states in the control group.

As with any policy-intervention analysis, there is a concern that the states that passed mandate-to-cover laws are systematically different from those that did not. A comparison of the demographic and labor force characteristics of women by legal environment is presented in Table II. Most of these measures will be considered as outcome variables in the results section. In 1982-83, before any states passed mandate-to-cover laws, women are comparable in terms of education, fertility timing, and the rate of multiple births, though birth rates are slightly higher for older women in mandate states. The labor force variables are also similar across states. [10]

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10. Hamilton and McManus [2005] perform a similar comparison of the ten states with any form of mandated insurance coverage for IVF to those without such a law. Using 1990 Census data, they find no significant differences in means for demographic categories such as education, labor

Also, with regard to possible endogeneity of the laws, it is important to note that lobbying efforts to promote such legislation generally argue that infertility is a life-altering disease which agents should be able to insure against through group insurance plans.[11] It does not seem to be the case that women support the laws so that they might be allowed to delay children. Therefore, any active delaying of fertility and subsequent changes in career outcomes should be viewed as an unintended and secondary effect. However, because concern about the endogeneity of the laws remains, and because Table II does show small differences across states, it is important to choose an econometric specification that is robust to these differences. In the estimation strategy described below, I include state fixed-effects and state-specific time trends to allow states to have their own trend, and any observed effects of the law should not be attributed to pre-existing conditions or trends.

B. Time-Series/Cross-Section Estimation

I will exploit this variation in the cost of infertility treatments using a differences-in-differences estimation strategy. Observations are at the individual level, and the key independent variable is the number of years that a woman's state of residence has had a strong mandate-to-

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force participation, income, and family size. They do find that law states were more likely to have had other similar insurance mandates. This is potentially a problem if states with mandates are more likely to require that insurance companies cover contraception. The first statewide contraceptive coverage law was passed in 1998, at the end of the data period used here. Results are robust to omitting data from 1999, which is the first year that contraceptive coverage might have affected women's fertility.

11. See for example, the advocacy groups Resolve [www.resolve.org] and The Inter-National Council on Infertility Information Dissemination, Inc. [www.inciid.org].

cover law. An increasing effect would be expected if women who have been covered for a longer period of time are more aware of the law and of infertility treatment options, or if they have had more time to adjust their behavior in response to the law. Models with a simple dummy variable for the law, leads and lags of the law, and non-linearities in years covered by the law were also considered and yielded qualitatively similar results.

It is important to notice that for the labor force participation and wage results, the timing issue is somewhat complicated. In order to strategically use infertility treatments to delay childbearing and for this to have important implications for their careers, women would have to know about the law when they were young enough for it to affect their fertility timing decisions. For example, being covered by a law for two years might affect the timing decision of a 25 year-old, but not a 45-year old. Therefore, in all tables I present specifications in which years covered by a mandate is interacted with age. I also create age-labor force participation and age-wage profiles by law coverage, with a cubic age trend.

To address concerns of law endogeneity described above, I allow for state fixed-effects, with year dummies. Even with state fixed-effects, however, there is still a problem of pre-existing trends. Infertility treatment use, for example, may have been increasing more rapidly in states that passed laws even before the laws were introduced. If these trends continued after the laws were passed, this would spuriously generate positive coefficients from differences-in-differences estimation. Therefore, the specification also includes quadratic state-specific time trends, and differences among states with different legal environments should not be attributed to pre-existing differences in trends. The equation estimated is:

(1) $$Y_{ist} = \alpha + \beta \bullet yearslaw_{ist} + \theta X_{ist} + \gamma_{is} + t_{is} + t_{is}^2 + \delta_{it} + \varepsilon_{ist}$$

where Y is the dependent variable, α is a constant term, X is a vector of individual

characteristics, γ represents state fixed effects, δ represents time dummies, and "yearslaw" is a variable indicating years of law coverage as defined above. The subscript i represents the individual, t represents the period, and s represents the state. For fertility regressions with a binary dependent variable, both a linear probability model and a probit model were estimated, and yielded very similar results. Because the interaction terms are more easily interpreted in the linear probability model, those results are discussed here. Finally, to address concerns about the size of tests based on differences-in-differences estimation, standard errors are clustered at the state level.[12]

V. Results

A. Multiple Births and Fertility Timing

Does empirical analysis reveal increases in multiple births and changes in mothers' ages in mandate-to-cover states that are consistent with the national trends described above? The results for fertility outcomes are presented in Tables III, IV, and V. First, I use the CDC's Natality Detail Files from 1982-1999 to consider multiple births, which should be interpreted as a proxy for infertility treatment use. The dependent variable is the number of children per birth,

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12. When there is positive serial correlation in the error term, conventional D-in-D estimates using OLS underestimate the standard errors. The relatively long time period of 18 periods used here increases the size of the bias. Therefore, any findings of an effect of insurance coverage laws using traditional OLS would be subject to criticisms regarding the size of the test. In fact, using simulations of a panel data set in which there are placebo interventions that have no effect by construction, Bertrand, Duflo, and Mullainathan [2003] find a significant effect using conventional D-in-D methods up to 45% of the time. The authors find that the clustering technique used here produces results with approximately the correct size.

where a singleton = 1, twins = 2, etc. Only white women are included, because women of different races have varying rates of multiple births. I examine women who are at least thirty because these women are most likely to be using infertility treatments as a result of age-related infertility (as opposed to chronic, early-onset infertility). Finally, I introduce controls for the age of the mother, mother's education level, birth order, and marital status. The controls for age are particularly important given that older women are naturally more likely to have multiple births.

In all but the last specification for regressions without the weak law states, there is a positive and statistically significant effect of the law on multiple births. The coefficient of 0.0025 in the fourth specification tells us that for each additional year of law coverage, the average number of children per birth increases by about 2.5 per 1,000, or by about 7.5 per 1,000 after five years. With a rate of 1,033 children per 1,000 births in the general population in 2001, this represents an increase of 22.7% in mandate states. Given rates of multiple births among women using and not using infertility treatments, we would need to see a 220.6% increase in infertility treatment use after a mandate was passed to get this increase in multiple births. While this figure seems large, it is comparable to effects found by Hamilton and McManus [2005] and is reasonable given the size of the cost reduction.[13] The interaction terms reveal that the effect of the mandate is slightly increasing in both age and education. This result is consistent with the fall in fertility with age and with the fact that more educated women are more likely to be

13. Hamilton and McManus [2005] find that states with a universal mandate had 41% more clinics than non-mandate states, and that cycles per clinic increased by 142% for women under 35 and by 193% for women over 35. The calculations used to create the 220.6% figure were based on a nationwide fertility therapy use rate of 1%, and a multiple birth rate of 3% for non-fertility therapy births and 37% for fertility therapies [CDC 2003].

covered by the mandate. Finally, notice that in the fifth specification, the positive and significant coefficient for the interaction of marital status tells us that the mandates affect the infertility treatment use of married women relatively more than single women.

Turning to fertility timing, Table IV shows the effect of the mandate on the percent of mothers between the ages of 40 and 49 with at least one child under six. This variable is constructed as a proxy for the birth rate for women ages 35-44, since we do not observe the precise ages of children in the CPS. The results in the table are the coefficients from a linear probability model, where the dependent variable is equal to one if the woman has a child under six and zero otherwise.

In the full specification without the weak law states, we see that each year of strong-law coverage results in a 0.36 percentage point increase in the probability that an older woman has a young child, or a 1.8 percentage point increase after five years of coverage. Given that 4.5% of college-educated women in this age group had a child under six in 1982-1983, this figure represents an increase of 40%--which is consistent with the 32% increase found by Schmidt [2005]. The effect is smaller for older women but larger for more educated and married women. If it is in fact career women who are using infertility treatments to delay children, the finding that more educated women are more responsive to the laws is not surprising. The result that married women are impacted more than single women is consistent with the findings for multiple births.

While these results are suggestive, the question still remains: Are women delaying children in anticipation of access to infertility treatments? It could simply be that women who tried to have children later in life and had difficulty conceiving were now able to do so with these technologies. To address this issue, I now consider birth rates among younger women, for whom birth rates should decrease if women are in fact delaying children. Table V presents these

results, for women ages 22-25 and ages 26-30. We do in fact see a decrease in each of the specifications of around 2.5 percentage points per year of coverage, which translates to about a 26% decrease in the birth rate to young women after five years of coverage. Thus, the hypothesis that some women are using infertility treatments to delay children finds support here.

B. Labor Force Participation Profiles

Do we see changes in women's labor market participation patterns in states with better access to infertility treatments? If fertility treatment does in fact allow women to delay childbearing, then we might expect to see increased labor force participation rates among women in their 20s and early 30s. Also, women may be working more during this period in order to qualify for employer-provided insurance and the infertility benefit. Because women who delayed childbirth would then have young children in their late thirties and forties, we might also expect to see decreased participation among women in this age group.

Here, the dependent variable of interest is the percent of women in the state who were in the labor force in the week before their March CPS interview.[14] Results from a probit regression including years covered by a law, a cubic age function, state fixed effects, year dummies, state-year specific time trends, and interactions of all variables with age are summarized in Table VI, and show a statistically significant effect of the laws. These coefficients are used to create the age-participation profiles shown in Figure V. Consistent with the above hypothesis, we do observe increases in labor force participation for covered women before age 35 and decreases after age 35. For 29 year olds, for example, the probability of participating in the labor force increases by 0.074 after five years of law coverage—which corresponds to a 10.86% increase

14. All labor force participation and wage data are from the March Current Population Surveys, 1982-1999.

above the national average for the rate before fertility treatments were introduced. For a 39 year old woman, on the other hand, the probability of participation decreases by 0.064 after five years of coverage, for a 9.39% decrease.[15] The age at which participation rates for women in mandate states dip below those of women in non-mandate states is slightly higher for married women, which may suggest further delay on the part of married women.

The results for fertility timing in the previous section suggested that better access to infertility treatments leads to delayed childbirth. Here, we see that career interruptions, which for women are often a result of childbirth, are delayed as well. The hypothesis that women with better access to infertility treatments will delay childbearing and career interruptions finds further support here.

C. Wage Profiles

The above changes in labor force participation rates suggest that women do delay career

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15. In regression results not reported here, an additional year of coverage increases the probability of labor force participation for the average woman under 35 by 0.0094, or by 0.047 after five years—which corresponds to a 6.9% increase above the national average before fertility treatments were introduced. The effect is in fact stronger for more educated women—the decrease in labor force participation after five years of law coverage is about one percentage point higher for a college-educated woman than for a woman with a high school degree. The coefficients on years covered for older women are negative and larger in magnitude than those for young women, though statistically insignificant. Compare this to Goldin and Katz [2002], who find that the growth in oral contraceptive use between 1970 and 1990 accounts for a 1.7-percentage point increase in the number of women entering professional occupations.

interruptions when they have the opportunity to do so. Does this option to delay then result in increases in wages for working women that are compatible both with the above model and the findings of Ellwood, Wilde, and Batchelder [2004]? Table VI also presents results for wages, where the estimation strategy is the same as with labor force participation. Here, the dependent variable is last year's average log hourly wage for women with at least a high school degree who worked 1,000 hours or more in that year. The sample is confined to women who work at least part time because we are interested in the effect of the timing of career interruptions on the wages of working women.

The laws are again found to have a statistically significant effect, and the profiles generated by the coefficients are presented in Figure V. Here, wages are lower for women in mandate states early in the career. For younger women, the mandates appear to have a negative effect on wages. A negative effect might be observed if employers pass the cost of the additional coverage along to women by reducing their wages (this is also true for older women), though the estimated effects are larger than those suggested by the findings of Griffin and Panak [1998]. The negative effect might also be due to selection bias. The results for labor force participation showed increases in rates for women under 35 in mandate states. If less-skilled women were pulled into the labor force, average wages for this group would decrease.

By the early forties, women in mandate states appear to experience higher wages on average, though these results could be affected by selection as well. Here, because older women in mandate states have lower participation rates, negative selection of dropouts could increase wages for working women. However, if more highly skilled women are the ones using infertility treatments and dropping out, the increase in wages could be understated. A full treatment of the selection issue is beyond the scope of this paper, and thus the results here should be interpreted

with caution. What can be said is that the results support a theoretical model in which the opportunity to delay childbirth results in higher wages for women in the mid-point of their careers and beyond, and greater human capital accumulation. Combined with the evidence of delayed career interruptions for these women seen earlier, it appears that improved access to infertility treatments might improve career outcomes for women.

D. Other Outcomes

Because access to infertility treatments appears to affect fertility decisions and career outcomes, one might also expect educational attainment and marriage to be affected. Using the same regression specification as in equation (1) in results not reported here, I found a statistically significant but practically small positive effect of the mandates on educational attainment. I also found small positive effects of the mandates on the probability of being married for educated women. This effect would be observed if women married to receive the mandated benefit, or if they delayed marriage leading to better matches. Unfortunately, timing of marriage is not observable in the CPS.

VI. Conclusion

This paper explores the effects that increased access to infertility treatments, in the form of a strong mandate-to-cover law at the state level, has on family and career outcomes. The expected first order effects of an increase in multiple births and in the presence of young children among women 40-49 are observed in states with a strong mandate-to-cover law. A decrease in births to young women suggests that at least some women are using infertility treatments to delay childbirth. I then show that the presence of a law is correlated with increased labor force participation for women ages 25-34, and decreased participation for women ages 35-44. The law is also associated with wage increases in the late thirties and early forties for working women,

which is consistent with wage-growth models featuring returns to experience and human capital accumulation.

This paper has not explicitly addressed the policy debate on the costs and benefits of mandated infertility insurance. However, it does speak to a potential benefit of mandated coverage that has been almost entirely overlooked in the past. As mentioned above, advocates of mandated coverage have traditionally focused on the argument that infertility is a life-altering disease which agents should be able to insure against through group insurance plans. The work here presents an additional and presumably unintended benefit of access to infertility treatments—improved labor market outcomes and an improved career/family tradeoff for women. Given that at least three laws have been recently introduced into the United States House of Representatives that would mandate some form of infertility insurance coverage nationwide, any further understanding of the costs and benefits of such a law is valuable.

In this paper, I have attempted to identify the effects of infertility treatments on the career-family tradeoff. Particularly in the case of labor market outcomes, the effects of widespread infertility treatment use have only recently started to emerge. Furthermore, improvements in the success of infertility treatments and further reductions in costs will certainly make them an even bigger part of women's fertility decisions in the years ahead. It is not hard to imagine generations of women in the near future who will seriously consider infertility treatments as an option when making their fertility timing decisions. An extended biological clock would likely affect not only women's birth and career outcomes, but also their marriage and education decisions and the health and well-being of mothers and their children. Therefore, the development of these technologies should provide fertile ground for labor and health economists in the coming years.

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Data Appendix

Missing data and data inconsistencies have been addressed in the following ways:

- The natality data did not differentiate between triplets and higher order births until 1989. Thus, for consistency, a new variable was created that topcodes births from 1989-1999 at 3, and this variable was used for the results presented here. In 1989, only 67 of the almost 4 million births involved quadruplets or higher, so the results should not be affected by the topcoding.
- Some states only report 50% of their births in earlier years (Arizona, California, Delaware, District of Columbia, Georgia, and North Dakota). For these states, the observations were doubled for use in individual-level regressions and before any sampling was done. For regressions at the state level, this was not necessary. By 1985, all states report 100% of their births.
- For results using CPS data, the final weight and the earnings weight were used.
- The labor force participation measure used here (last week's labor status) is not available from the CPS in 1994, so that year is dropped from the participation and wage analysis.
- The wage regressions include women who worked 1000 or more hours in the previous year, with hourly earnings between $1 and $200. The hours requirement reduces the sample of women with at least 12 years of education between the ages of 25 and 44 from 179,693 to 108,554. The use of the CPS earnings weight further reduces the sample reported in the wage regressions to 52,630.

Table I
Infertility Insurance Coverage by State and Strength of Law, 1982-1999

State	Year Passed	Mandate to Cover--Strong	Mandate to Cover--Weak	Mandate to Offer
Maryland	1985	X		
Arkansas	1987	X		
Hawaii	1987	X		
Massachusetts	1987	X		
Montana	1987		X	
Texas	1987			X
California	1989			X
Connecticut	1989			X
Rhode Island	1989	X		
New York	1990		X	
Illinois	1991	X		
Ohio	1991		X	
West Virginia	1995		X	

States with a strong mandate-to-cover law have laws that cover at least 35% of insured women and must cover in vitro fertilization. Weak mandate-to-cover laws either cover less than 20% of insured women or had significant loopholes and also excluded IVF coverage. Mandate to offer laws require that insurance companies offer a policy to employers, but employers may choose not to purchase these plans. A few states adopted laws and New York and Connecticut strengthened their laws after 1999, which is beyond the period studied here. Source: Resolve [2004], National Conference of State Legislatures [2004].

Table II
Characteristics of Women in 1982-1983, By Future Legal Environment

Variable	Pass a Strong Law (6 States)	Do Not Pass a Strong Law (45 States)	Do Not Pass Any Law (38 States)
Average Education, All Women	12.5220 (0.0342)	12.5103 (0.0114)	12.5139 (0.0136)
Average Education, Non-mothers	12.6968 (0.0561)	12.7821 (0.0189)	12.6930 (0.0236)
Average Age of Mothers, 1st birth	23.2781 (0.0096)	22.8981 (0.0033)	22.7307 (0.0040)
Children per Birth	1.0213 (0.0002)	1.0206 (0.0001)	1.0207 (0.0001)
% Women Age 40-49 with Child Under 6	0.0562 (0.0063)	0.0441 (0.0018)	0.0409 (0.0021)
% Women Age 22-25 with a Child	0.4670 (0.0162)	0.4773 (0.0054)	0.4875 (0.0068)
% Women Age 26-30 with a Child	0.6101 (0.0148)	0.6474 (0.0047)	0.6583 (0.0058)
% in Labor Force Last Week, Women Ages 25-34	0.7045 (0.0100)	0.6823 (0.0033)	0.6860 (0.0041)
% in Labor Force Last Week, Women Ages 35-44	0.6885 (0.012)	0.6829 (0.004)	0.6913 (0.005)
Log Hourly Wages Last Year, Women Ages 25-34	1.7310 (0.027)	1.7916 (0.009)	1.7576 (0.012)
Log Hourly Wages Last Year, Women Ages 35-44	1.8509 (0.0363)	1.7983 (0.0106)	1.7724 (0.0126)

Standard errors are in parenthesis. States that pass insurance laws after 1999 are classified as not having a law. A strong law is defined as a law that covers 35% of women in the state and covers in-vitro fertilization. Data obtained from CDC Natality Detail Files 1982-1983, March CPS 1982-1983.

Table III
Effect of Being Covered by a Strong Mandate-to-Cover Law on Multiple Births

Independent Variable	Sample Without Weak Law States					All States
Years Mandate has Existed (Yrslaw)	0.0035**	0.0029**	0.0024**	0.0025**	0.0010	0.0010
	(0.0007)	(0.0008)	(0.0008)	(0.0009)	(0.0013)	(0.0012)
Age		0.0008**	0.0008**	0.0001	0.0001	0.0004**
		(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
Age*Yrslaw		0.0002**	0.0002**	0.0002**	0.0002**	0.0001**
		(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Education			0.0013**	0.0021**	0.0021**	0.0021**
			(0.0002)	(0.0002)	(0.0002)	(0.0002)
Education*Yrslaw			0.0001**	0.0001	0.0001	0.0001
			(0.0001)	(0.0001)	(0.0001)	(0.0001)
Birth Order				0.0055**	0.0055**	0.0052**
				(0.0003)	(0.0003)	(0.0003)
Birth Order*Yrslaw				-0.0001	-0.0001	0.0001
				(0.0001)	(0.0001)	(0.0001)
Married					0.0025**	0.0032**
					(0.0010)	(0.0009)
Married*Yrslaw					0.0016**	0.0015**
					(0.0005)	(0.0005)
Constant	1.0233**	1.0203**	1.0162**	1.0002**	0.9979**	0.9972**
	(0.0009)	(0.0010)	(0.0014)	(0.0017)	(0.0024)	(0.0019)
Observations	1,027,490	1,027,490	1,027,490	1,027,490	1,027,490	1,441,603
R-squared	0.0021	0.0023	0.0025	0.0042	0.0042	0.0040

* Significant at 10% level, ** significant at 5%. Results are the coefficients from an OLS regression that includes state fixed effects, year dummies, and state-year specific time trends. The dependent variable describes the number of children per birth (singleton=1, twins=2, etc). The variable age is constructed as age - 29, and education is years of education - 11. The sample is a 10% sample of all birth certificates in the CDC Natality Detail Files 1982-1999. The sample is limited to white women over thirty with at least 12 years of education.

Table IV
Effect of Being Covered by a Strong Mandate-to-Cover Law on Presence of Young Children, Women 40-49

Independent Variable	Sample Without Weak Law States		All States
Years Mandate has Existed (Yrslaw)	0.0031**	0.0036**	0.0032**
	(0.0011)	(0.0013)	(0.0012)
Age		-0.0138**	-0.0144**
		(0.0006)	(0.0006)
Age*Yrslaw		-0.0006**	-0.0005**
		(0.0003)	(0.0003)
Education		0.0083**	0.0079**
		(0.0005)	(0.0004)
Education*Yrslaw		0.0006**	0.0006**
		(0.0002)	(0.0002)
Married		0.0401**	0.0411**
		(0.0027)	(0.0026)
Married*Yrslaw		0.0019**	0.0018*
		(0.0009)	(0.0009)
Black		0.0226**	0.0240**
		(0.0039)	(0.0046)
Black*Yrslaw		-0.0023	-0.0024
		(0.0015)	(0.0015)
Other		0.0170**	0.0189**
		(0.0066)	(0.0031)
Other*Yrslaw		-0.0005	-0.0007
		(0.0026)	(0.0025)
Constant	0.0422**	0.0626**	0.0648**
	(0.0035)	(0.0045)	(0.0032)
Observations	107,411	107,411	148,201
R-squared	0.0050	0.0482	0.0484

* Significant at 10% level, ** significant at 5%. Results are the coefficients from a linear probability model that includes state fixed effects, year dummies, and state- specific time trends. The dependent variable is equal to 1 if the woman has a child under the age of 6 and 0 otherwise. The variable age is constructed as age - 39, and education is years of education - 11. Data are from the March CPS 1982-1999, and March CPS weights are used. The sample was limited to women age 40-49 with at least 12 years of education.

Table V
Effect of Being Covered by a Strong Mandate-to-Cover Law on Births to Young Women

	Women 22-25			Women 26-30		
	Sample Without Weak Law States		All States	Sample Without Weak Law States		All States
Independent Variable						
Years Mandate has Existed (Yrslaw)	-0.0302	-0.0250*	-0.0271*	-0.0219**	-0.0239**	-0.0241**
	(0.0215)	(0.0136)	(0.0138)	(0.0088)	(0.0068)	(0.0064)
Age		0.0243**	0.0237**		0.0361**	0.0341**
		(0.0033)	(0.0026)		(0.0012)	(0.0012)
Age*Yrslaw		-0.0003	-0.0002		-0.0006	-0.0004
		(0.0017)	(0.0018)		(0.0004)	(0.0004)
Education		-0.0837**	-0.0821**		-0.0753**	-0.0756**
		(0.0032)	(0.0035)		(0.0017)	(0.0014)
Education*Yrslaw		0.0009	0.0007		-0.0001	0.0000
		(0.0014)	(0.0014)		(0.0009)	(0.0009)
Married		0.2206**	0.2270**		0.3471**	0.3538**
		(0.0087)	(0.0087)		(0.0060)	(0.0052)
Married*Yrslaw		0.0038	0.0031		0.0003	-0.0004
		(0.0031)	(0.0032)		(0.0026)	(0.0026)
Black		0.2256**	0.2120**		0.1983**	0.1854**
		(0.0078)	(0.0089)		(0.0107)	(0.0095)
Black*Yrslaw		-0.0030	-0.0013		0.0010	0.0025
		(0.0066)	(0.0067)		(0.0018)	(0.0016)
Other		0.0545**	0.0460**		0.0415**	0.0175*
		(0.0186)	(0.0123)		(0.0121)	(0.0101)
Other*Yrslaw		-0.0112**	-0.0101**		-0.0008	0.0019
		(0.0050)	(0.0048)		(0.0050)	(0.0048)
Constant	0.4442**	0.4333**	0.4228**	0.6311**	0.3048**	0.3152**
	(0.0085)	(0.0100)	(0.0105)	(0.0088)	(0.0133)	(0.0103)
Observations	51,281	51,281	71,621	70,647	70,647	98,445
R-squared	0.0165	0.1874	0.1819	0.0200	0.2435	0.2470

* Significant at 10% level, ** significant at 5%. Results are the coefficients from a linear probability model that includes state fixed effects, year dummies, and state- specific time trends. The dependent variable is equal to 1 if the woman has a child and 0 otherwise. The variable age is constructed as age - 21, and education is years of education - 11. Data are from the March CPS 1982-1999, and March CPS weights are used. The sample was limited to women in the specified age group with at least 12 years of education.

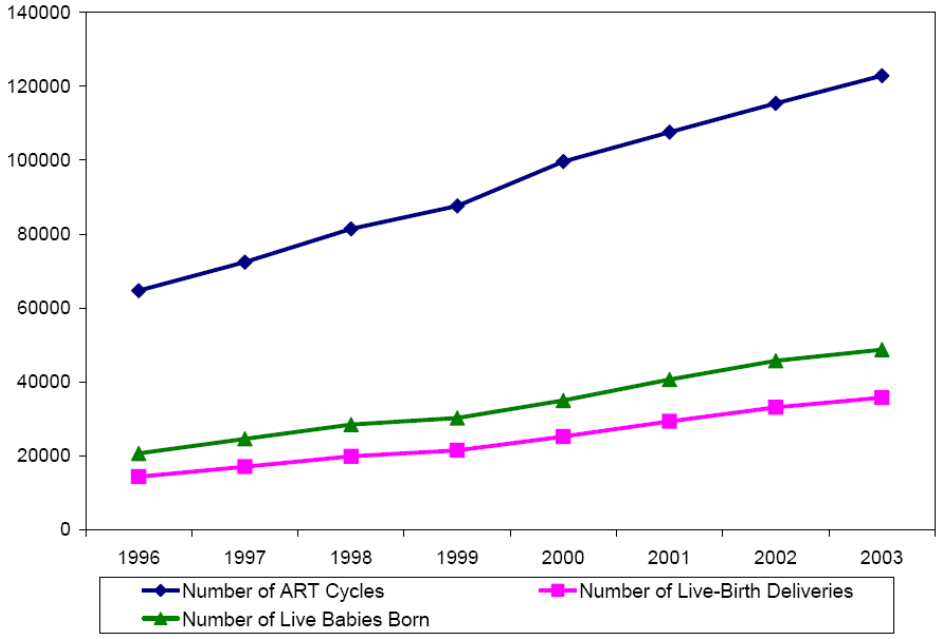
Table VI
Age-Labor Force Participation and Age-Wage Profiles

	Dependent Variable	
	In Labor Force Last Week	Log Hourly Wage Last Year
Independent Variable	Marginal Effects from Probit Regression	Coefficients from OLS Regression
Years Covered by Mandate (Yrslaw)	0.0253** (0.0076)	-0.1011* (0.0537)
Age	-0.0095** (0.0018)	0.03770** (0.0048)
Age^2	0.0011** (0.0002)	-0.0021** (0.0006)
Age^3	-0.0000** (0.0000)	0.0000** (0.0000)
Yrslaw*Age	-0.0022* (0.0012)	0 .0076** (0.0031)
Yrslaw*Age^2	-0.0001 (0.0001)	-0.0001 (0.0002)
Yrslaw*Age^3	0.0000 (0.0000)	0.0000 (0.0000)
Education	0.0337** (0.0018)	0.0866** (0.0031)
Education*Yrslaw	0.0001 (0.0005)	0.0005 (0.0008)
Education*Yrslaw*Age	0.0000 (0.0000)	-0.0000 (0.0001)
Married	-0.1373** (0.0057)	0.0538** (0.0113)
Married*Yrslaw	0.0020 (0.0022)	0.0013 (0.0039)
Married*Yrslaw*Age	0.0003** (0.0001)	0.0003 (0.0002)
Constant		1.4890** (0.0206)
Observations	254,788	39,065
Chi-squared/F-statistic for Yrslaw and all Yrslaw Interactions	424.56	7.74
Pseudo R-squared/R-squared	0.0420	0.2968

* Significant at 10%, ** significant at 5%. Results in the first column are the marginal effects from a probit regression with last week's labor force participation as the dependent variable, and results in the second column are the coefficients from an OLS regression with log hourly wage as the dependent variable. Both regressions include state fixed

effects, year dummies, state-year specific time trends, race dummies, interactions of race dummies with yrslaw, and interactions of all variables with age. The variable age is constructed as age-24, and education is years of education - 11. Data are from the March CPS 1982-1999, and CPS weights are used. States with weak mandates are excluded. The sample is limited to the specified age group with at least 12 years education. These results are used to create Figure IV and Figure V.

Figure I: Number of ART Cycles Performed, Number of Live-Birth Deliveries, and Number of Live Babies Born Using ART, 1996-2003



The number of live babies exceeds the number of deliveries because of multiple births (twins, triplets, etc.). Source: Center for Disease Control, Assisted Reproductive Technology Success Rates [2003]. Figure is a reproduction of Figure 43 from this report.

Figure II: Average Number of Babies per First Birth, White Women by Age

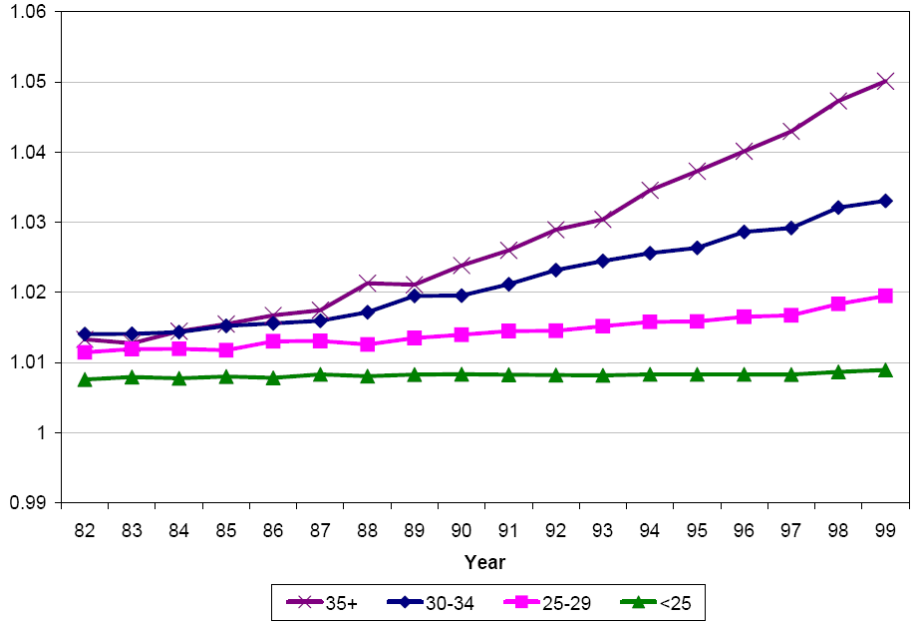
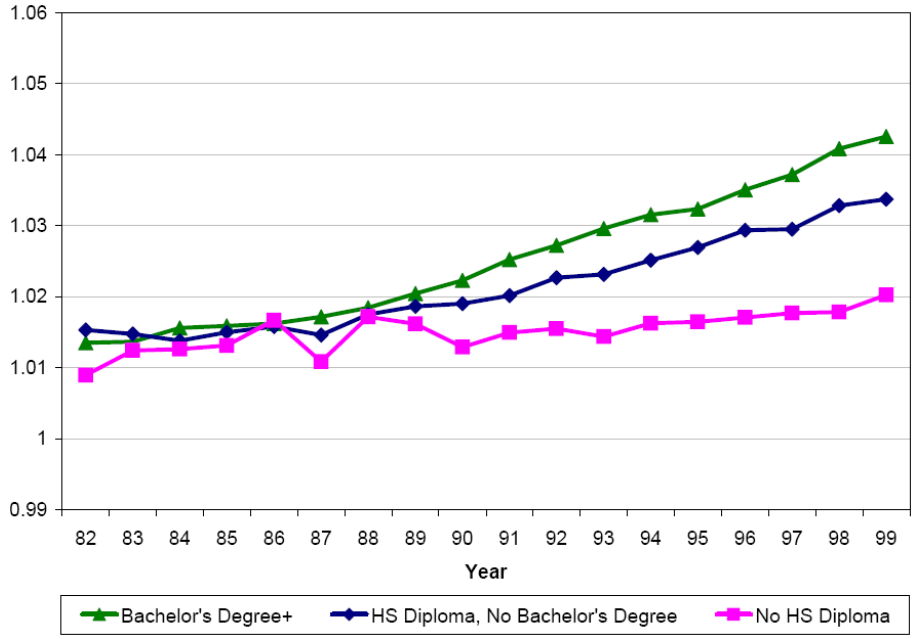


Figure III: Average Number of Babies per First Birth, White Women, Age 30+, by Education Level



Figures II and III: Average number of babies per first birth is calculated by counting a singleton birth as 1, a twin birth as 2, a triplet birth as 3, etc. Source: Center for Disease Control, Natality Detail Files [1982-1999].

Figure IV
Predicted Labor Force Participation, by Age and Years Covered by Mandate

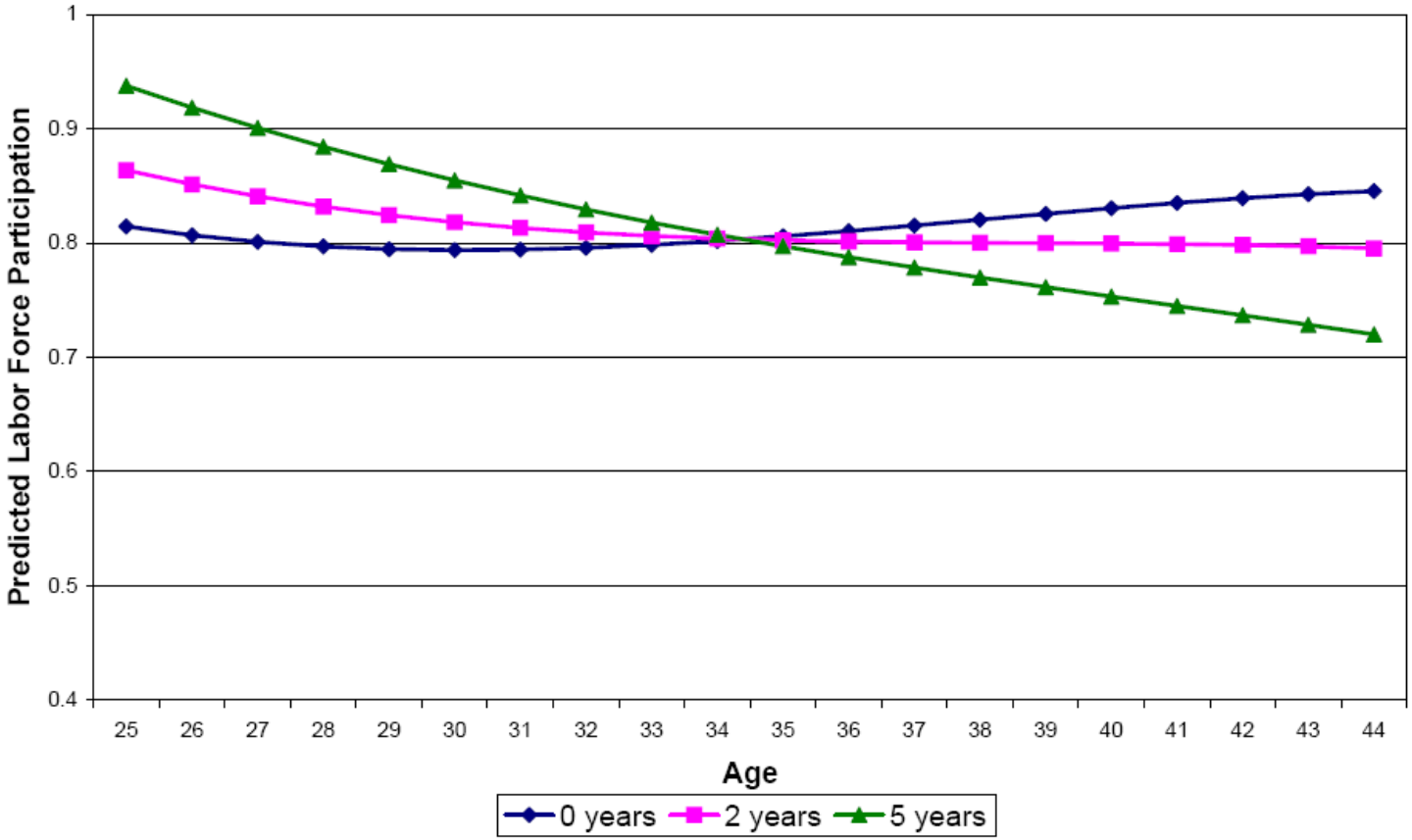


Figure is created from results presented in Table VI. Values for education and marital status are equal to the population means at 13.68 and .67 respectively, with race = white and year = 1999. Labor force participation is based on responses for the previous week. Data are from the CPS March Files, and March CPS weights are used.

Figure V
Predicted Hourly Wage, by Age and Years Covered by Mandate

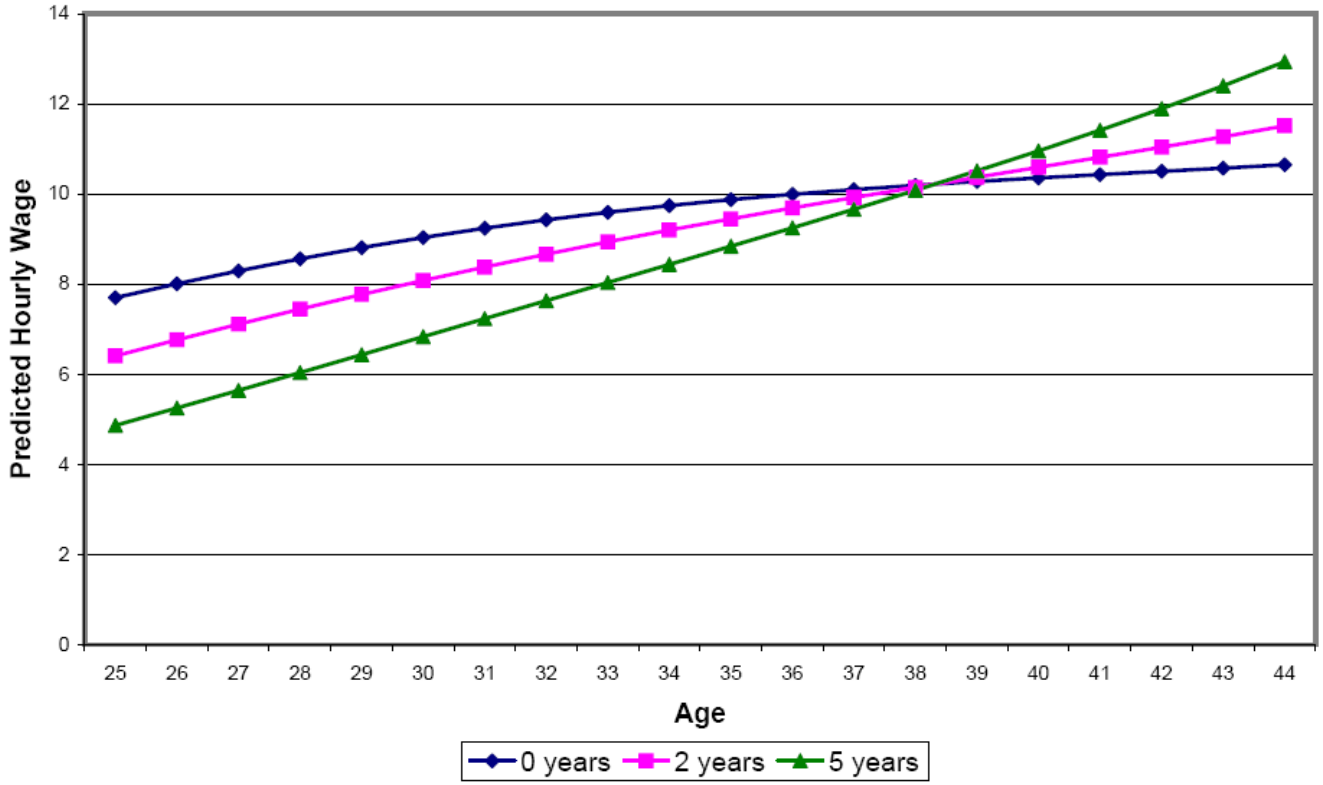


Figure is created from results presented in Table VI. Values for education and marital status are equal to the population means at 13.68 and .67 respectively, with race = white and year = 1999. Log hourly wage is calculated from the previous year's income and hours worked, and the sample is limited to women who worked 1000 hours or more, with hourly earnings between $1 and $200. Data are from the CPS March Files, and March CPS earnings weights are used.