Health Insurance, Fertility, and the Wantedness of Pregnancies: Evidence from Massachusetts

Maria Apostolova
Department of Economics
University of Kentucky

Aaron Yelowitz
Department of Economics
University of Kentucky

November, 2013

Abstract: Health insurance reform in Massachusetts lowered the cost of pregnancy and the cost of preventing pregnancy (through increased access to reliable contraception). We examine fertility responses for women of childbearing age, and on net, find no effect of increasing health insurance coverage. This masks substantial heterogeneity, however. For married women aged 20-34 – who have high latent fertility and for whom pregnancies are typically wanted – fertility increased by approximately 1 percent. For unmarried women in the same age range – for whom pregnancies are typically unwanted – fertility declined by 9 percent. For other age/marital status groups, there was very little fertility response, in part because of low latent fertility or minimal gains in insurance coverage. Pregnancy wantedness increased in the aggregate through a combination of increasing wanted births and decreasing unwanted births.

Keywords: Health Insurance, Fertility, Moral Hazard, Pregnancy wantedness
JEL Classification: I13, I18, J13

Contact information: Apostolova: Department of Economics, Gatton School of Business and Economics, 550 South Limestone Street, Lexington, KY 40506; Telephone (859) 257-7646; Email address: maria.apostolova@uky.edu, Website: www.mariaapostolova.com. Yelowitz: Department of Economics, Gatton School of Business and Economics, 550 South Limestone Street, Lexington, KY 40506; Telephone (859) 257-7634. Email address: aaron@uky.edu, Website: www.yelowitz.com.

We are grateful to Frank Scott, Chris Bollinger, Jenny Minier, Jim Ziliak, Glen Mays, Beth Munnich and Gordon Dahl for helpful comments. Apostolova acknowledges the generous dissertation support from the Horowitz Foundation for Social Policy.
I. Introduction

The Patient Protection and Affordable Care Act (PPACA) is the first successful attempt in the U.S. to provide near-universal health insurance coverage at the national level but in recent years similar policies have been implemented at the state and local levels.\(^1\) Among these reforms, the Massachusetts health care law is the most prominent example of increasing health insurance coverage. The Massachusetts experience has been used to study questions of critical importance like changes in health outcomes and costs following the expansion of health insurance but given that the Massachusetts legislation served as a model for the design of the PPACA, the answers to these questions have broader implications about the future consequences of the national reform.

It is well established that coverage rates increased as a result of the reform in Massachusetts, although there is disagreement about the exact magnitude of the effect.\(^2\) One of the intended consequences was that the out-of-pocket cost of expensive medical events, which include pregnancy-related expenses, was reduced due to increased insurance coverage.\(^3\) Thus, women of child-bearing age who wanted to have children could have been incentivized by the reform to plan and carry out a pregnancy if they were previously uninsured. In addition

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\(^1\) In 2003 Maine enacted the “Dirigo Health Reform”, in 2006 Vermont adopted “Catamount Health”, in 2007 the city of San Francisco launched the program “Healthy San Francisco”, and in 2006 Massachusetts enacted the “Act Providing Access to Affordable, Quality, Accountable Health Care”. The first three reforms rely primarily on subsidies for purchasing health insurance, while the Massachusetts law also includes an individual mandate, an employer mandate, and an expansion in Medicaid.

\(^2\) Official estimates for the uninsured rate in 2008 were 2.6 percent, but Yelowitz and Cannon (2010) find that uninsured rates are underreported because the reform incentivizes people to hide their true insurance status if they are uninsured. According to their estimates, which exclude imputed values for health insurance questions, the uninsured rate is 3.8 percent.

to moral hazard from expanding health insurance coverage, adverse selection may still be important in the design of the Massachusetts legislation and the PPACA. Even with an individual mandate, individuals are given the option to remain uninsured by paying a penalty (which starts off relatively small, and increases over time) or purchase less comprehensive coverage. \(^4\) With guaranteed issue and modified community rating, women of childbearing age might then purchase more comprehensive coverage when anticipating pregnancy and childbirth and avoid such coverage if they do not anticipate pregnancy. \(^5\) The ease of movement in and out of generous plans due to adverse selection further subsidizes high-cost, anticipated medical events like pregnancy.

In addition to lowering the costs of having a baby, the Massachusetts law also lowered the costs of preventing a pregnancy because the use of family planning services and access to reliable contraception became easier. \(^6\) As a result, women of child-bearing age who did not want to get pregnant might have increased their use of reliable birth control and decreased their fertility. In terms of the PPACA, funding for birth control, family planning, and abortion has become a very controversial topic with several attempts to alter the relevant provisions of the

\(^4\) For example, catastrophic health insurance plans offered under the PPACA are available to those under age 30. It covers essential health benefits (including maternity and newborn care) but the high deductibles (approximately $6300 for an individual) would likely deter many women who anticipate pregnancy from purchasing such a policy. See https://www.healthcare.gov/can-i-buy-a-catastrophic-plan/ for eligibility requirements (accessed 11/19/2013), and https://www.nebraskablue.com/~/media/Broker%20Communications/2014_Individual_Renewal/92139.pdf for an example of specific provisions related to deductibles and pregnancy expenses.

\(^5\) Feldstein (2013) argues such a design will encourage those who are healthy to strategically remain uninsured until they have a potentially costly medical diagnosis. See http://www.project-syndicate.org/commentary/martin-feldstein-on-how-america-s-health-care-reform-could-unravel (accessed 11/19/2013).

\(^6\) The most reliable, non-permanent, forms of contraception include IUDs and implants (with less than 1 pregnancy per 100 women each year), and shots, pills, rings and patches (with 2-9 pregnancies per 100 women each year). Each of these requires either a doctor’s prescription or contact with a health care provider and can involve significant out-of-pocket costs if uninsured. Less reliable forms of contraception (with between 15-25 pregnancies per 100 women) include diaphragms, male condoms, female condoms, withdrawal, sponges, cervical caps, and spermicide. These typically do not require contact with physicians or health care providers, and often entail lower out-of-pockets costs for the uninsured. See http://www.plannedparenthood.org/health-topics/birth-control/birth-control-effectiveness-chart-22710.htm, (accessed 11/19/2013).
law. At present, the use of federal funding for abortion, except in rare cases, is prohibited by an executive order of the President. The national legislation also includes a mandate for contraception coverage applicable to all health insurance. In addition, all new health plans must cover certain women’s preventive services with no co-payments, including the full range of FDA-approved contraception methods and contraceptive counseling.⁷ To the extent that this mandate remains unchanged, studying the response of birth rates in Massachusetts can provide useful insights into the broader demographic effects of the PPACA.⁸

In this paper we use the exogenous nature of the Massachusetts healthcare reform to identify the effect of changes in insurance coverage on fertility behavior. Previous studies exploring the fertility effects of Medicaid expansions for the poor have found limited evidence of moral hazard effects. To our knowledge, our study is the first to examine fertility responses in the context of the Massachusetts healthcare reform. We rely on the American Community Survey (ACS), which is uniquely positioned to study this issue because of its explicit questions on fertility and sizable samples. When we use a straightforward difference-in-differences strategy, we do not find any change in fertility rates for the full sample of women of childbearing age (15-44 years old), a result which is consistent with some earlier work. Since insurance coverage rates also vary based on socioeconomic characteristics (rather than just by state and year), we further parameterize the changes in coverage. Our preferred specification

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replaces the standard difference-in-differences estimator with this more parameterized version of insurance coverage. Even with this parameterized specification, we do not find an effect on fertility when we examine all women or stratify the sample based on age. Our key finding emerges when we analyze the fertility effect by age and marital status – the Massachusetts law increased fertility rates for married women aged 20-34 by 1 percent and decreased fertility rates for unmarried women in the same age group by 9 percent. These opposite-signed results potentially reflect the differential degree of wantedness of pregnancies that varies by marital status. These effects cancel each other out which is why we do not observe a fertility effect in the aggregate. Fertility rates for women aged 15-19 and 35-44 did not change as a result of the Massachusetts reform. The lack of effect is not surprising given that younger women did not experience large gains in insurance coverage (hence, identification is more difficult) and fertility rates for older women are generally very low (hence, there is a heterogeneous behavioral response). These results are robust to the inclusion of different sets of control variables and a variety of specification checks.

The remainder of the paper is arranged as follows. In Section II we survey the existing literature on the Massachusetts health insurance reform and the fertility responses to expanding public health insurance coverage. Section III provides a description and timeline for health insurance reform in Massachusetts. In Section IV we discuss the predicted effects of expanding health insurance coverage on fertility, and show how the response should vary by observable characteristics. Section V describes our data. Section VI presents the empirical framework and the findings are shown in section VII. Finally, Section VIII concludes.
II. Literature Review

Our paper adds to an emerging literature evaluating the causal effect of the Massachusetts health care reform on various outcomes of significant policy relevance. This is an important topic because of the unique opportunity to draw conclusions and make predictions about the possible effects of the Affordable Care Act, the main provisions of which are closely aligned with those of the Massachusetts health care reform. Insurance coverage and health care utilization are among the main outcomes that have been examined to date by the literature. Our paper adds to the broader understanding of the effects of this policy by evaluating the extent to which an important and permanent outcome – individual fertility behavior – changes in response to increased health insurance coverage.

Several studies examine the effect of this reform on changes in insurance status since a key goal of the Massachusetts health insurance law was to achieve nearly-universal coverage. There is a general consensus that uninsured rates decreased although there is some disagreement on the exact magnitude. Early studies by Long (2008) and Long, Stockley, and Yemane (2009) find reductions in the uninsured rate of approximately 50 percent among all adults aged 18-64, while Kolstad and Kowalski (2012a) attribute only a 36 percent decrease in

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9 Other outcomes of interest are health outcomes, crowd-out of private health insurance, labor market effects, and adverse selection. Courtemanche and Zapata (2012) find that the reform lead to improved self-reported health for minority, low-income, and near-elderly individuals. Miller (2012b) finds evidence that the probability for a child to be in excellent health increased by 6 percentage points after implementation. Yelowitz and Cannon (2010) show reductions in self-reported excellent health but also show gains in good self-reported health. Long (2008) finds that the share of people who had employer-provided health insurance remained the same in the early stages of the Massachusetts reform and Miller (2012b) demonstrates that among children private insurance coverage increased by 8-10 percentage points. Kolstad and Kowalski (2012a) find limited evidence that private health insurance was replaced by Medicaid among hospitalized patients only, while Yelowitz and Cannon (2010) report that private coverage decreased among children and low-income adults. Kolstad and Kowalski (2012b) examine the efficiency and welfare impact of the reform in the labor market. Hackman, Kolstrad, and Kowalski (2012) find that increased coverage resulted in lower costs per enrollee which is evidence for the existence of adverse selection in the pre-reform Massachusetts health insurance market.
the uninsured rates to the new law.\textsuperscript{10, 11} Yelowitz and Cannon (2010) show that Massachusetts residents with incomes between 150-300 percent of the Federal Poverty Level (FPL), who are most likely to be affected by the reform, tend to underreport their insurance status in the CPS, which likely results in an underestimation of the officially reported uninsured rate.

The reform-induced gains in health insurance coverage were different based on socioeconomic characteristics because the law did not impact all state residents in the same way. Several studies find that the drop in uninsured rates was more substantial among low-income adults and young adults while it was very modest for older and wealthier individuals.\textsuperscript{12} For example, Niedzwiecki (2013) shows that insurance coverage increased mostly for poor adults and individuals aged 19-30 years. On the other hand, the reform did not significantly affect children’s health insurance status. Long, Stockley, and Yemane (2009) find that uninsured rates among children did not change immediately after the reform while Miller (2012b) shows that children and young teens (less than 18 years old) saw an economically insignificant increase in insurance coverage in the order of 2 percentage points.\textsuperscript{13} These findings are not surprising given that coverage rates of children and teenagers are typically quite high because they are often eligible under a parent’s health insurance policy or through the Medicaid.

\textsuperscript{11} Long (2008) finds that the uninsured rate decreased from 13.0 percent from fall 2006 to 7.1 percent in fall 2007 among adults. Long, Stockley, and Yemane (2009) find that the uninsured rate decreased from 12 percent to 5.4 percent in the first year after the reform.
\textsuperscript{12} Long (2008) shows that the overall reduction in uninsured rates was driven by a 46 percent drop for adults with incomes up to 300 percent of the FPL. She also finds a statistically significant decrease in uninsured rates for adults with incomes over 300 percent of the FPL but Long, Stockley, and Yemane (2009) do not find such an effect using CPS data.
\textsuperscript{13} Even though the overall effect on insured rate among children is small, Miller (2012b) does find that the reform altered substantially the type of insurance coverage children were getting.
In addition to lowering uninsured rates, the Massachusetts reform also changed healthcare utilization; the results suggest greater efficiency afterwards. The use of preventive healthcare services increased following the implementation of the law. Miller (2012a) examines how increased health insurance coverage affected emergency room (ER) use and finds that the reform resulted in reductions between 5-8.4 percent in ER visits mainly for nonemergency medical events and during normal business hours. Kolstad and Kowalski (2012a) observe a similar effect on emergency room hospitalizations which decreased by 5.2 percentage points following the reform and were largely driven by reductions for patients from poor areas. They also examine preventable admissions and find reductions for severe patients of about 2.7 percent. Niedzwiecki (2013) shows that the use of preventive healthcare services increased through a reduction in avoidable hospitalizations, although he also finds an overall increase in emergency department visits.

To our knowledge, there are no studies that examine moral hazard in the context of the Massachusetts reform as it relates to individual fertility behavior. Most work focusing on fertility-related moral hazard effects uses the exogenous variation in state Medicaid expansions that occurred in the 1980s and 1990s. A notable exception is Leibowitz (1990) who uses data from the RAND health insurance experiment to show that pregnancy rates and births increased temporarily by 29 percent in response to random assignment of free health insurance coverage.14

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14 Insurance coverage was randomly assigned and varied based on the level of co-insurance rate and the maximum out-of-pocket expenditure.
Work using the Medicaid expansions mostly shows a heterogeneous response of birth rates based on socio-demographic characteristics with little evidence for an overall effect. \(^{15}\) Joyce, Kaestner, and Kwan (1998) examine birth rates of unmarried young women (19-27 years old) with less than 12 years of education and find that Medicaid expansions increased birth rates of white women by 5 percent. They further demonstrate that the fertility response varies by income because birth rates increased more during the first wave of the Medicaid expansions when coverage was extended to individuals with incomes up to 100 percent of the FPL. \(^{16}\) Zavodny and Bitler (2010) analyze the fertility effects of Medicaid expansions for all women of childbearing age (15 to 44 years old). They measure the fertility effect of Medicaid expansions by using the fraction of women eligible or the statutory Medicaid income eligibility threshold. They find no evidence for an overall effect, but do find that a 100 percentage point increase in the income eligibility threshold for Medicaid would increase birth rates by 7.7 percent among low-educated white women. This is consistent with the idea that there are heterogeneous effects for different subgroups in the population. DeLeire, Lopoo, and Simon (2011) construct birth rates by state, year, quarter, and demographic cell based on age, marital status, and education for women aged 15-44. Recognizing that Medicaid expansions affected women differently, they construct a policy variable as the fraction of the population eligible for Medicaid which varies by quarter, state, year, and demographic cell. They find weak evidence that expansions of public health insurance coverage had an effect on birth rates. In models that only include state, year, and quarter fixed effects, and state-year fixed effects, but do not

\(^{15}\) An exception is Yelowitz (1994) who does find an overall increase in birth rates of 5 percent for women aged 15-44. He also finds a differential effect by race and marital status.

\(^{16}\) The second wave of Medicaid expansions in the late 1980s and early 1990s qualified pregnant women for Medicaid if family income was less than 185 percent FPL.
control for demographic characteristics, increasing Medicaid eligibility has a positive and statistically significant effect on birth rates of approximately 1.2 percentage points for whites and 2.4 percentage points for African-Americans. However, this result is quite sensitive to changes in the specification. A limitation of this study is that it ignores the potentially differential effect of insurance on birth rates for groups with different demographic characteristics, which is a key contribution of our study.

The increased availability of health insurance not only lowers the cost of having a baby but also lowers the cost of not having a baby, because almost all health plans cover contraception (and some cover abortion). Importantly, in Massachusetts after the reform, the publicly subsidized “Commonwealth Care” plan covered a full range of family planning services including abortion care. Using non-representative data collected through surveys and interviews Dennis et al. (2012) conclude that access to affordable contraception improved for low-income women post-reform even though they faced several challenges. Family planning community centers have provided critical support for overcoming these obstacles (Dennis et al, 2009) and might play an important role of entry to the health care system for particular groups of women in the context of the national health care law (Gold, 2009). An important distinction between the Massachusetts experience and the PPACA is that national health care reform

17 The PPACA contraception mandate requiring all health insurance plans to offer contraceptive coverage at no cost exempts churches and non-profit religious groups from this provision while their female employees can still have contraceptive coverage paid directly by the insurance company. This mandate also spurred litigation on the basis of employers’ religious beliefs where for-profit business owners also fight for an exemption. This issue is pending review by the Supreme Court in November 2013. Source: http://www.nytimes.com/2013/11/02/us/court-rules-contraception-mandate-infringes-on-religious-freedom.html?ref=contraception accessed on 11/17/2013.
18 Particular problems identified in this study for low-income women include understanding and maintaining health insurance coverage, filling out prescriptions, and obtaining appointments with a healthcare provider mostly stemming from the fact that before the reform these women’s contraceptive needs were met almost exclusively by family planning community centers but after the reform these women had to, independently, understand their coverage, select a doctor, make an appointment, and fill a prescription.
mandates contraception coverage at no cost while in Massachusetts there was still an element of cost-sharing.\textsuperscript{19} Since both the cost of having a baby and the cost of not having a baby are typically reduced by increased availability of health insurance, the net effect on birth rates is ambiguous.

The literature examining the effect of Medicaid expansions on fertility behavior focuses primarily on net birth rates with few studies analyzing the impact on abortion rates.\textsuperscript{20,21} Zavodny and Bitler (2010) do not find an overall effect on abortion rates but due to data limitations they are unable to stratify the sample by demographic characteristics. Joyce, Kaestner, and Kwan (1998) examine birth rates and abortion rates separately in order to split the confounding effects of subsidized health insurance. They also do not find an overall effect but show that restricted Medicaid funding for abortions led to a decrease in abortion rates of approximately 10 percent among white women, which is consistent with increased use of family planning services for this group. Joyce and Kaestner (1996) also show a reduced likelihood of abortion for young, low-educated nonblack women. Kearney and Levine (2009) specifically examine the impact of income-based expanded Medicaid eligibility for family planning services on teen and adult birth rates and contraceptive use.\textsuperscript{22} They find that fertility of women in different age groups responds differently to the income-based expansions with

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\textsuperscript{20} Since 1972 Medicaid has covered the use of contraception (Kearney and Levine, 2009).

\textsuperscript{21} The Hyde Amendment, passed in 1976, prohibits the use of federal funds for abortions except in cases of rape and incest or when the life of the mother is endangered. As a federally-funded program, Medicaid does not cover abortions under both fee-for-service and managed care frameworks (National Committee for a Human Life Amendment, “The Hyde Amendment”, April 2008 from: http://nchla.org/issues.asp?ID=1, from: accessed on 11/16/2013).

\textsuperscript{22} Kearney and Levine (2009) also analyze the effect on abortions and sexual activity.
\end{footnotesize}
the largest effect among 18-24 year-olds where birth rates declined by 5.1-6.8 percentage points. They are able to show that this result is mainly driven by an increased probability of contraceptive use. These findings provide strong motivation for analyzing the fertility responses separately for different demographic groups because of the variation in their underlying fertility rates and possibly wantedness of the pregnancy. We expand on this hypothesis in the next section.

Our survey of the relevant literature indicates that the Massachusetts reform achieved its goal of expanding insurance coverage although some groups saw larger gains in insurance coverage than others. It also led to increased health care services utilization. Studies focusing on the Medicaid expansions generally find that the fertility behavior of different demographic groups is affected differently with little evidence for an overall effect. Policies expanding subsidized family planning services have been found to be effective in reducing unintended births through an increased use of contraception.

III. The Massachusetts Health Care Reform Timeline

The sweeping healthcare law dramatically changed the landscape of the Massachusetts health insurance market. The implementation of the reform began in October 2006 and continued through July 2007 (see Table 1 for a timeline of its major stages). During that transition period, the state expanded coverage under Medicaid and the Children’s Health Insurance Program (CHIP) for children with family incomes up to 300 percent FPL and raised

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enrollment caps for adults. The law also allowed adults younger than 26 years to remain on their parents’ health insurance plans. In addition, the state provided subsidized coverage to individuals with family incomes up to 300 percent FPL on a sliding-scale basis with full subsidies for those with incomes up to 150 percent FPL (Commonwealth Care).

The individual mandate – effective July 2007 – required individuals to purchase health insurance or pay a fine equal to their personal state income tax exemption (in the first year) or a penalty of up to 50 percent of the lowest health insurance premium they would be eligible for (after the first year).

Standards of coverage and costs were established for insurance companies who had to offer health insurance at (modified) community-rated premiums, mandated minimum coverage, and maximum premiums irrespective of preexisting health conditions and claims made. The employer mandate required employers with 11 or more full-time employees to offer employer-provided health insurance or face modest penalties.

As demonstrated by the existing literature, the reform resulted in expanding health insurance to the vast majority of Massachusetts residents. However, the gains in insurance were not evenly distributed among the whole population with some demographic groups gaining more from than others. For example, low-income individuals with family income close to the FPL were more likely to be covered by Medicaid so the reform did little to improve coverage rates for them. Likewise, wealthier individuals who tend to get insurance from employers did not gain much. We use these uneven gains to predict fertility responses in the next section.

IV. Predicted Effects of Expanding Health Insurance on Fertility
Predicting the effects of health insurance reform on fertility in the aggregate is not a straightforward exercise because in addition to the availability of health insurance, several other elements, like latent fertility (proxied by age) and wantedness of children (proxied by marital status) factor into the decision to have a baby. Younger women have higher fertility rates than older women, because older women are both more likely to have reached their desired fertility and more likely to suffer from infertility. All else equal, married women are more likely to become pregnant and carry a baby to term than unmarried women, because pregnancies are more likely to be planned and wanted.

These considerations suggest that if expansion of health insurance coverage has effects on fertility, the effect would vary by age and marital status. All else equal, expanding insurance coverage will increase fertility of married women (wanted pregnancies) due to lowering the out-of-pocket cost of pregnancy, while reducing fertility for single women (unwanted pregnancies) due to better access to reliable contraception. As health insurance becomes more widely available, births for younger women should increase more than for older women due to the higher latent fertility rates of the former group.

In addition to the considerations related to latent fertility and wantedness of pregnancies, the insurance coverage gains varied widely within age/marital status group. There were clearly gains in Massachusetts (relative to the rest of New England) from 2008 onward (relative to 2003-2006). Moreover, those gains in Massachusetts after the reform varied by family income. Some groups experienced minimal gains in insurance (such relatively affluent

25 The majority of children born as the result of an unplanned pregnancy are born to women who are either single or cohabiting. See http://www.thenationalcampaign.org/resources/pdf/FactSheet-Consequences.pdf, accessed 11/20/2013.
women who were often covered by private insurance), while others experienced much larger
gains (such as near-poor women). Larger gains in coverage should lead to larger fertility
responses within each age-marital status cell. Figure 1 summarizes the expected fertility effects
by age, marital status, and income which reflect variations in latent fertility, child wantedness,
and insurance gains. A woman who was young, married, and near-poor experienced larger
relative gains in insurance coverage and would have been relatively more likely to have a baby
than a similar woman who was richer. A woman who was young, unmarried, and near-poor
experienced larger relative gains in insurance coverage and would have seen larger relative
reductions in fertility compared to a similar woman who was richer.

It is also important to note the importance of the interaction of age and marital status.
One would expect that the fertility responses for older women – regardless of whether
pregnancies were wanted or insurance gains were large – would be much smaller due to lower
latent fertility. For very young women (teenagers), one might expect smaller fertility responses
as well if for no other reason than the insurance gains were typically much smaller.

V. Data Description: ACS, CPS and Vital Statistics

Our primary data source is the Census Bureau’s American Community Survey (ACS)
Public Use Microdata Sample (PUMS). We use the one-year sample of the ACS PUMS for the
years 2003-2011 (excluding the transition year of 2007 when the reform was being phased in).
Starting with the 2005 PUMS, approximately one percent of all households in the U.S. were
surveyed (in 2003 and 2004, the samples are approximately 40 percent the size of subsequent
years). As a consequence, we are able to examine the fertility responses in Massachusetts
relative to other New England states, while retaining sizable samples. Moreover, we are able to examine responses for narrow demographic groups, such as married women aged 20-34, where we can more accurately characterize the wantedness of pregnancies and latent fertility. The ACS often asks similar questions to the now phased-out decennial Census long forms. Unlike most other household surveys that the Census Bureau conducts, respondents are required by law to participate in the ACS.26

Important for our purposes, the ACS directly asks fertility questions pertaining to each woman aged 15-50 in a surveyed household. Specifically, the survey asks “Has this person given birth to any children in the past 12 months?” Other datasets, such as the March Current Population Survey (CPS) Annual Social and Economic Supplement (ASEC) do not directly ask about fertility; instead, one might impute fertility from the presence of infants in a household (i.e. 0-year-olds on the household roster). It is clear that such an imputation strategy might have difficulty in assigning an infant to a mother if there is more than one woman of childbearing age in the household. Perhaps more importantly, our investigation of the ACS shows that many infants are potentially not living with their mothers, and this non-presence is related to socioeconomic circumstances. As Table 2 shows, across all years, only 80 percent of households where a birth was reported had an infant (0 years) present. Some of this is certainly confusion about the survey question, however, because the fraction of households reporting a birth who also have a zero- or one-year-old present is 88 percent.27 Nonetheless, there is still

26 Source: http://www.census.gov/acs/www/Downloads/language_brochures/ACSQandA_ENG10.pdf. See Title 13, United States Code, Sections 141, 193, and 221. The decennial Census is a notable exception in that it is mandatory.

27 If a household misinterpreted “the last 12 months” with “the last year” or “the last calendar year”, they might report a one-year-old as a birth.
an important difference between births and presence of very young children. Roughly 10 percent of infants live in a household where there is not a woman reporting a birth.

In Table 3, we show that the modest disconnect between reported births and presence of infants is largely a function of socioeconomic circumstances. We examine 242,006 women aged 15 to 44 in the 2003-2011 ACS across the entire U.S. who reported a birth (and where that woman was the only one in the household reporting a birth). The outcome of interest is whether an infant (defined as age 0) is missing on the household roster. Unmarried, non-white, and less-educated women are far more likely – 8 to 12 percentage points – to not have a baby present in the household; this may be unsurprising if the father lives in a separate household or if members of the extended family, like grandparents, take care of the child. More surprisingly, the likelihood of missing infants increases sharply with age; 35-39 year-olds are 6 percentage points more likely to not have an infant present, while 40-44 year-olds are nearly 24 percentage points more likely to not have an infant present. We interpret these age results somewhat differently than the socioeconomic results, however. Fertility is quite low among these age groups – especially 40-44 year olds – and many of the affirmative responses to the fertility question in the ACS could simply be survey errors for these women. Given this possibility, we also do our empirical analysis separately by age group.

There is one unfortunate drawback of the ACS, however. The ACS did not start asking questions on health insurance status until 2008, which is the beginning of our “post” period. As a consequence, we rely on the Current Population Survey (CPS) to derive insurance rates, and append these to each woman in the ACS sample based on various socioeconomic characteristics, state of residence, and time period. We use the 2004-2012 CPS March
Supplements, which cover calendar years 2003-2011 (excluding 2007), to compute insured rates for women by demographic category, region and time.\textsuperscript{28,29} We do this separately for Massachusetts and the other five New England states combined (Connecticut, Vermont, Maine, Rhode Island, and New Hampshire). The demographic categories are based on age, income, and marital status. There are 6 age groups (15-19, 20-24, 25-29, 30-34, 40-44), 4 income groups (<150% FPL, 150-250% FPL, 250-300% FPL, >300% FPL), 2 marital statuses (married and unmarried). For each demographic group, we create coverage rates for 2 regions (Massachusetts and the rest of New England) and 2 periods (“before” period including calendar years 2003-2006 and “after” period including calendar years 2008-2011).\textsuperscript{30} The total number of insurance coverage cells is therefore 192 (6 ages x 4 income x 2 marital x 2 regions x 2 periods).

We define a woman as “uninsured” if she is not covered by private health insurance, Medicare, Medicaid, or CHAMPUS. The insurance coverage rate is then the weighted ratio of the number of insured women in each cell to the total number of women in the cell.

Insurance coverage rates were highest among teenage women (15-19 year-olds) and older women (aged 35-44) both in Massachusetts and the rest of New England in 2003-2006 (Figure 2a). This is expected because teenagers are typically covered under their parents’ health

\textsuperscript{28} It is thought that CPS answers to health insurance questions are a blend of current coverage and coverage in the previous year. Swartz (1986) argues that CPS respondents ignore the precise wording of the health insurance questions, and instead answer the question as if it referred to coverage as of the survey date.

\textsuperscript{29} A similar approach was used by DeLeire, Lopoo, and Simon (2011) where the policy variable is an index of Medicaid eligibility that varies by quarter, state, year, and demographic cell, which is based on age, race, marital status, and education.

\textsuperscript{30} We follow the existing literature in treating 2006 as a “before” year because the earliest provisions went into effect in October 2006. See Hackmann, Kolstad, and Kowalski (2012), Long, Stockley, and Yemane (2009) and Yelowitz and Cannon (2010), all of whom use annual data. Given the time horizon for pregnancy, and the wording of the question in the ACS, the vast majority of pregnancies in this year would have been prior to reform. In addition, the ACS respondents take the survey throughout the year (and it is not possible for us to identify the date when the survey was answered). Virtually all studies classify 2007 – where the individual mandate is implemented in the middle of the year – as a transition year. Our interest lies in the effects of the fully phased-in reform; thus we focus on 2008 onward as the “after” period.
insurance plan or Medicaid and older adults are more likely to be insured due to improved economic circumstances. The age groups with lowest coverage rates were 20-24 and 25-29 year-olds because young adults leaving college were often no longer covered on a parent’s plan and may not have a job with fringe benefits like health insurance. The gains in insurance coverage in Massachusetts following the reform were therefore most pronounced for these age groups (Figures 2b and 2c); coverage increased by almost 14 percentage points for 20-24 year-olds and 8 percentage points for 25-29 year-olds. The changes among teens and older adults were quite modest in comparison. In contrast to Massachusetts, the rest of New England experienced relatively small gains and even reductions in coverage rates for some age groups.

Figures 3a, 3b, and 3c show that coverage rates are higher for married women than unmarried women, because of the availability of spousal health insurance coverage. Massachusetts’ reform had an equalizing effect for unmarried women; insurance coverage increased by almost 8 percentage points.

The changes in coverage rates by income are illustrated in Figures 4a, 4b and 4c. Insurance coverage was initially highest for women with incomes over 300 percent of the FPL, and the coverage gains were very small (2 percentage points). The coverage gains were somewhat limited for the poorest women with incomes less than 150 percent of FPL because many had health insurance coverage through Medicaid. In contrast, the middle group with incomes between 150-300 percent FPL saw increases in insurance coverage of 12-14 percentage points.

Finally, although women between 15 and 44 are often categorized as being of “child-bearing” age, birth rates vary tremendously by age group. Older women in the sample are more
likely to have reached their desired number of children, and as such, one may not expect the same fertility response to insurance coverage that would see with a younger woman. We calculate a latent fertility variable that represents the propensity of a woman to give birth that varies by age, marital status, and race. To construct latent fertility, we use two datasets because the variable is computed as a ratio of the number of births occurring to women within a demographic cell divided by the total number of women in that group and there is no single dataset that provides this information. For the numerator we use the Center for Disease Control and Prevention’s Vital Statistics data which records all births in the U.S. for a particular year. We use natality data from 2003, the first year of our analysis, to establish baseline fertility rates. We calculate the number of births for six age groups (16-19, 20-24, 25-29, 30-34, 35-39, and 40-44), two marital statuses (married and unmarried), and two races (white and non-white) for a total of 24 demographic cells. For the denominator we use weighted ACS data from 2003 to obtain the total number of women at risk for birth by demographic cell.

As an example, to compute the latent fertility variable for married non-white women aged 20-24 years, we divide the number of births to women in this demographic cell (59,984) by the total number of women in the U.S. within the same cell (431,350) and obtain a latent fertility rate of 13.9 percent for non-white, married women aged 20-24 years. The inverted-U shape of latent fertility in Figure 5a illustrates wide variations of the propensity for having a baby, with young women aged 20-34 being most likely to give birth. Birth rates among married women are significantly higher for each age group than for unmarried women (Figure 5b), reflecting a higher degree of child wantedness. Conditional on age and marital status, race is not an important factor affecting latent fertility. We use the constructed fertility rates in two
ways. First, the obvious variation of latent fertility provides strong motivation for stratifying the sample, both by age and by age and marital status. Second, as part of the robustness checks, we interact the latent fertility rate with the insured rates to evaluate how the change in insurance coverage rates affect the birth rates for different levels of latent fertility.

VI. Empirical Framework

As is well recognized, the Massachusetts reform creates a quasi-experiment to evaluate the impact of expanding health insurance coverage on various outcomes. The natural starting point for our examination of fertility is a straightforward difference-in-differences (DD) estimator estimated from a linear probability model:

\[
BIRTH_{ijt} = \beta_0 + \beta_1 MASS_j \star POST_t + \beta_2 MASS_j + \beta_3 POST_t + \beta_4 X_{ijt} + \varepsilon_{ijt}
\]

where \(BIRTH_{ijt}\) is a dummy variable equal to one if woman \(i\) in state \(j\) at time period \(t\) had a child in the past 12 months, \(MASS_j\) is a dummy variable for living in Massachusetts (relative to the other New England states – Connecticut, Maine, New Hampshire, Rhode Island and Vermont), and \(POST_t\) is a dummy variable for the years 2008 and beyond (relative to the years 2003-2006).\(^{31,32}\) We also include controls for the woman’s education (high school dropout, high school graduate with college graduate being omitted), whether the woman has changed residence in the past year, whether she has served in the military and whether she is a

---

\(^{31}\) Results from a probit model are similar. 
\(^{32}\) We omit the transition year of 2007.
non-citizen. In later specifications, we also interact age, marital status, and poverty levels. The coefficient estimate on $\beta_1$ is then interpreted as the DD estimator.\footnote{The large majority of papers studying the effect of the Massachusetts health care law use some form of difference-in-differences identification strategy. See for example, Kolstad and Kowalski (2012a), Courtemanche and Zapata (2012), Yelowitz and Cannon (2010), Long et al. (2009), and Miller (2012a).}

Although transparent, there are many reasons to go beyond the specification in equation (1). Most importantly, although the near-universal health reform in Massachusetts leveled coverage rates across groups, as discussed in the previous section, there were very different gains based on one’s initial socioeconomic circumstances. Thus, we create a parameterized version of equation (1) by attaching to each woman the insurance coverage rate based on her state, time period, and demographic group (48 categories, 6 age x 4 income x 2 marital status).\footnote{Similar methods for constructing a policy variable are consistently used by the literature examining the effect of Medicaid expansions on various outcomes. This measure is typically the fraction of the population eligible for Medicaid (DeLeire et al., 2011; Zavodny and Bitler, 2010; Currie and Gruber, 2001).} Thus, equation (2), which forms our baseline specification of insurance gains on fertility, is:

$$(2) \quad BIRTH_{ijt} = \beta_0 + \beta_1 INSURED_{djt} + \beta_2 DEMOG_d + \beta_3 X_{ijt} + \delta_s + \delta_t + \epsilon_{ijt}$$

where $INSURED_{djt}$ is the fraction of demographic group $d$ covered in region $j$ in period $t$.\footnote{Since the variation in INSURED is at a higher level than the individual, all standard errors are clusters at the DEMOG*STATE*YEAR level. The significance of the results is very similar if we simply cluster at the STATE*YEAR level and the significance is much stronger if we cluster at the STATE level alone. Thus, we view our results at the most conservative approach.} It is likely that the key components of $INSURED_{djt}$ – especially demographics like age and marital status – have a direct effect on fertility; thus, we include a full set of dummy variables for demographic group ($DEMOG_d$), as well as state and year fixed effects ($\delta_s$ and $\delta_t$). The estimate of the impact of insurance coverage, $\beta_1$, is identified from how Massachusetts’ changing health insurance landscape over time interacted with different demographic groups. Since the
identification of the insurance effect comes from the interaction of state, time and demographics, we present further specifications that show the conclusions are relatively robust to including finer controls like state-year interactions.

One key drawback to running equation (2) on the full sample, however, is that such a specification imposes an equal marginal impact on fertility for gains in insurance coverage. There are clearly reasons to think this should not be the case.\textsuperscript{36} As discussed in the prior section, older women who are of childbearing age are likely to have reached their desired number of children; as a consequence, one might not expect much impact on fertility for them. Moreover, gains in insurance coverage not only reduce the cost of having a baby, but also reduce the cost of preventing a pregnancy. One would expect that pregnancies are much more likely to be wanted for married women, and unwanted for single women. Thus, the estimate from equation (2) above could combine both positive and negative fertility responses. As a consequence, in addition to examining the full sample, we separately stratify by age group, and also age group and marital status.\textsuperscript{37} Further, we provide confirmation that latent fertility matters greatly for the actual fertility response by estimating equation (3):

$$BIRTH_{idjt} = \beta_0 + \beta_1 INSURED_{djt} \times FERT_d + \beta_2 INSURED_{djt} + \beta_3 DEMOG_d + \beta_4 X_{ijt} + \delta_s + \delta_t + \epsilon_{ijt}$$

\textsuperscript{36} Joyce, Kaestner, and Kwan (1998) only include young, single, and low-educated women in their sample.
\textsuperscript{37} Zavodny and Bitler (2010) stratify the sample by race-marital status and race/education to analyze the fertility effects.
where $FERT_d$ is the latent fertility rate estimated from the 2003 Vital Statistics and ACS data by age/marital status/race groups, and the other variables are defined as before.38 Assuming that pregnancies are wanted rather than unwanted, then one would expect $\beta_1$ to be positive.

VII. Empirical Results

We first summarize fertility rates and insurance rates for our sample. Overall, the sample consists of more than 500,000 women aged 15-44 in Massachusetts and surrounding states. Nearly 8 percent report a birth in the past year, across all years. In addition, our imputed insurance rate is nearly 92 percent – reflecting both the changes in Massachusetts after 2007, and the high overall level of coverage in New England. Consistent with the Vital Statistics data, fertility rates vary dramatically by woman’s age. More than 12 percent of women aged 20-34 report having a baby in the previous 12 months, approximately five times the rate of women aged 15-19 or 35-44. The fertility differences are especially pronounced by marital status; nearly 21 percent of married women aged 20-34 reported having a baby, more than three times the rate of unmarried women in the same age group.

Our first attempt at estimating the impact of insurance coverage on fertility is shown in Table 4, corresponding to the difference-in-differences specification in equation (1). For both the full sample, as well as each age group, one would conclude the expansions in insurance had little effect on fertility. In all cases, the coefficient estimate is substantively small and insignificant. As noted, however, this specification ignores many important aspects about the fertility decision and the Massachusetts reform, in particular, the uneven gains in insurance

38 The DEMOG variable is then changed to be the interaction of age, marital status, income and race. The FERT variable is not included as a separate regressor because its variation is subsumed by the DEMOG variable.
coverage, the different latent fertility rates by age group, and the differential wantedness of pregnancies between married and unmarried women.

Thus, we turn to Table 5, which estimates equation (2), by including the parameterized insurance rate. As in the previous table, when one looks at the full sample or particular age groups, insurance gains appear to have no effect on overall fertility. Yet, as shown in columns (3) and (4), there are significant and opposite signed effects for unmarried and married women aged 20-34. Recall this is the group with the highest latent fertility rate. Although not shown, coefficient estimates are insignificant and much smaller for other age/marital status groups. For unmarried women aged 20-34, insurance coverage increased by 11.4 percentage points due to the Massachusetts law. With a coefficient of -0.0459, this would imply that fertility fell by -0.52 percentage points. Since the pre-reform baseline fertility in the ACS was 5.99 percent among this group, then fertility fell by 8.7 percent. For married women, in contrast, gains in insurance coverage led to increased fertility. The overall gain in insurance coverage was much more modest – 2.4 percentage points – which leads to an increase in fertility of 0.23 percentage points from a much higher baseline of 20.11 percent. Thus, among married women, fertility increased by around 1 percent.

Given the striking differences in insurance coverage on married and single women, one may ask whether the marriage decision itself is endogenous to the law. Yelowitz (1998) found that the expansions in Medicaid in the 1980s and 1990s led to higher marriage rates. The key difference between the Medicaid expansions and the more recent Massachusetts context is that Medicaid had been traditionally targeted to poor female headed families on cash welfare

---

39 We ran difference-in-difference estimates similar to equation (1) to get the change in insurance coverage.
(poor married families were largely ineligible); hence the expansions in Medicaid opened up eligibility to married couples and on the margin, created incentives to get married. The expansion in Massachusetts, on the other hand, applied to all groups and was essentially neutral with respect to marriage. In any case, we have run difference-in-differences regressions (parallel to equation 1) where the outcome is whether the woman is married; for both the full sample as well as each age group, the estimate is insignificant and substantively small in all cases. 40

One concern about Table 5 is that the INSURED variable is a complicated function of demographics, state and year. Although we include dummy variables for each of the main effects, a concern may be that there are interactions of DEMOG, STATE and YEAR that are also correlated with fertility decisions independent of the expansions in health insurance coverage. For example, the Great Recession may have affected income or employment in Massachusetts differently than the rest of New England, and those differences, rather than health insurance coverage, may be driving the fertility decisions. In Table 6, we test the sensitivity of the coefficient estimates to these kinds of stories for women aged 20-34. For comparison, Columns (1) and (5) replicate the specification and findings for unmarried and married women in the prior table. By including STATE*YEAR effects in columns (2) and (6), we see that such state-specific shocks have little impact on the underlying conclusions (if anything, the negative impact is slightly stronger for unmarried women). The next columns add DEMOG*YEAR interactions (in addition to the STATE*YEAR interactions). Although no longer significant, the coefficient

40 Chen (2013) examines “marriage lock” and finds reductions in divorce and increases in marriage due to the Massachusetts reform. She uses New Jersey and sometimes Connecticut as control states. Our findings on marriage rates come from examining a younger set of women, and using New England states (which does not include New Jersey) as our control group. Further investigation is certainly needed to reconcile her results on marriage with our non-results.
estimate for single women remains quite similar. For married women, the fertility effect is still positive, significant, and much the same magnitude as before. Finally, when one fully saturates the model, by including fixed effects for STATE*YEAR, DEMOG*YEAR and DEMOG*STATE (thus, the identification comes only from the interaction of STATE*YEAR*DEMOG in the INSURED variable), then neither estimate is significant, although it is reassuring that the actual estimate is quite similar to the baseline estimate. Overall, the findings for both unmarried women and married women hold up remarkably well to including additional controls.

In addition to exploring the sensitivity of the specification to additional controls, we consider one other important issue. There may be a concern that the generous health insurance benefits in Massachusetts – with community-rated premiums and guaranteed issue – makes the state a more attractive place for individuals with high expected medical costs, such as pregnant women and therefore encourage migration. It is important to note that if this is the case, one might expect to see increases in fertility for all age groups, rather than increases for married women and decreases for unmarried women. Nonetheless, selective migration is clearly a concern.41 Using the CPS, Yelowitz and Cannon (2010) find that in-migration in Massachusetts fell relative to other New England states as a result of the law, and this effect was particularly pronounced among adults aged 18-29. They interpret the drop in in-migration for the young – who are largely healthy and have low expected medical costs – as arising from the greater implicit tax on them arising from community rating and individual mandates. The

41 Gelbach (2004) shows that among women likely to use welfare, movers move to higher-benefit states. Aizer, Currie and Moretti (2007) and Marton, Yelowitz and Talbert (2012) explicitly account for the possibility that relatively attractive Medicaid health insurance packages might induce migration across counties within a state.
same factors that generate implicit taxes for most young adults also create implicit subsidies for pregnant women.

The ACS asks about one-year-migration patterns, and allows us to test this hypothesis. We restrict the sample to women who did not move across state lines in the previous year, and estimate equation (2) on each group. Our results are quite similar to the baseline results in Table 5. For the 94 percent of unmarried women aged 20-34 who did not move across state lines, the coefficient estimate (standard error) is now -0.0422 (0.0222) compared with the initial estimate of -0.0459 (0.0216). For the 95 percent of married women aged 20-34 who did not move across state lines, the coefficient estimate (standard error) is now 0.0821 (0.0538) compared with the initial estimate of 0.0982 (0.0504). In both cases, the small fraction of cross-state moves would appear to have greater responses to insurance coverage relative to non-movers, but the basic conclusions remain unchanged by restricting the sample to non-movers.

Finally, equation (3) interacts latent fertility with gains in insurance coverage and includes the main effect for insurance coverage. In all specifications in Table 7, the interaction term is positive, suggesting that gains in insurance coverage have positive effects on fertility for those with higher latent fertility. The main effect of insurance coverage is negative. This would be consistent with the findings in earlier tables, since married women have higher fertility rates (and thus, the interaction term is relatively important), while unmarried women have lower fertility rates (and thus, the main effect is relatively important).

Finally, we have examined an alternative definition of fertility. Recall that earlier tables found that some women reported pregnancies but did not have infants living in the household, and on many dimensions the missing infants varied logically with socioeconomic characteristics
like marital status, race and education. We have run similar specifications to our baseline result, but where childbirth is now defined as having an infant present on the household roster. Such a measure has difficulty in linking the infant to a mother when there are multiple women of childbearing age in a household, or when the mother is missing from the household. Although we continue to find reductions in fertility for unmarried women aged 20-34 and increases in fertility for married women aged 20-34, the coefficient estimates are roughly one-third as large and not statistically significant. In our view, this provides further evidence of the value of the self-reported pregnancy question over an approach that imputes past pregnancy based on household configurations.

In summary, although the expansions in health insurance coverage had close to zero net effect on fertility in Massachusetts, substantial heterogeneity exists for different demographic groups. Our findings suggest that latent fertility and the wantedness of children, along with differential gains in coverage, help explain opposite signed effects for married and unmarried women aged 20-34, and also explain the non-existence of effects for other groups. Married women in this age bracket increase their fertility when experiencing gains in coverage because pregnancies are largely wanted and underlying fertility is high. Single women, on the other hand, decrease their fertility because pregnancies are largely unwanted and better access to contraception helps them prevent pregnancy. For women aged 35 and older, latent fertility is quite low (and insurance coverage was typically high prior to the reform), so the overall fertility responses are small (and insignificant). For teenagers, fertility rates are also quite low, many pregnancies are unwanted, and insurance coverage was fairly high prior to reform. Thus, we find small and insignificant effects for them too.
VIII. Conclusions and Discussion

In this paper we examine the effect of the Massachusetts health care reform on a woman’s probability of having a baby. Although we find zero net effect on fertility for women aged 20-34, this masks substantial heterogeneity across married and unmarried women (which proxies for child wantedness). Among young married women, fertility increased by 1 percent while fertility decreased by 9 percent for young unmarried women. We find no effect on birth rates for teens or older women either in total or when we stratify by marital status.

Whether the reform simply shifted the timing of births or changed the total number of children a woman will have in her lifetime remains an open question. Data over a longer period than four years needed to assess the long-term fertility effect of the Massachusetts reform. Regardless of whether the reforms simply reflect a change in timing, the proportion of unintended pregnancies (those that are mistimed, unplanned or unwanted) fell as a result of the law.

Our results are informative about the future consequences of the PPACA which was largely modeled after the Massachusetts law. Expanding insurance would likely wanted pregnancies on a national scale. To the extent that the contraception-related provisions of the legislation remain unchanged, the national reform might also lead to a decrease in unwanted births. Such a reduction in unwanted births could have favorable implications for long-term crime trends in much the same way that legalizing abortion did (Donohue and Levitt, 2001). Preventing an unwanted birth might lead to increased investment in women’s own human capital and the human capital of their children thus increasing the overall level of future human capital and affecting the future rates of economic growth. To the extent that the reduction in
unwanted births is due to increased use of contraception (rather than increased abortion rates), potential savings could be realized in terms of resources spent on unintended pregnancies.
IX. References


Table 1
Timeline of Health Care Reform Implementation

<table>
<thead>
<tr>
<th>Date</th>
<th>Event</th>
</tr>
</thead>
<tbody>
<tr>
<td>April 2006</td>
<td>Health Care Reform legislation passed</td>
</tr>
<tr>
<td>July 2006</td>
<td>Federal Government approves Medicaid waiver for health care reform</td>
</tr>
<tr>
<td>October 2006</td>
<td>Plan Type I for Commonwealth Care open for enrollment (for residents at 100% of FPL)</td>
</tr>
<tr>
<td>January 2007</td>
<td>Plan Types II, III and IV for Commonwealth Care open for enrollment (for residents between 100% and 300% of FPL)</td>
</tr>
<tr>
<td>March 2007</td>
<td>Deadline for Connector Board to set minimum “creditable” coverage standards</td>
</tr>
<tr>
<td>May 2007</td>
<td>Commonwealth Choice plans available – individuals and small businesses can buy insurance</td>
</tr>
<tr>
<td>July 1, 2007</td>
<td>Individual mandate to purchase health insurance</td>
</tr>
<tr>
<td></td>
<td>Deadline for employers to provide health insurance to full-time employees</td>
</tr>
<tr>
<td></td>
<td>Deadline for merging the individual and small-group insurance markets</td>
</tr>
<tr>
<td>January 2008</td>
<td>Individual mandate penalty: 50% of premium per month if uninsured</td>
</tr>
</tbody>
</table>

## Table 2

**Household ages and the ACS fertility question**

“*Has this person given birth to any children in the past 12 months?*”

<table>
<thead>
<tr>
<th></th>
<th>Households where a birth is reported</th>
<th>Households where a birth is not reported</th>
<th>Fraction of zero-year olds in household reporting birth</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>All Years</strong></td>
<td>79.87%</td>
<td>0.20%</td>
<td>92.52%</td>
</tr>
<tr>
<td>2011</td>
<td>76.89%</td>
<td>0.17%</td>
<td>92.56%</td>
</tr>
<tr>
<td>2010</td>
<td>78.70%</td>
<td>0.17%</td>
<td>93.24%</td>
</tr>
<tr>
<td>2009</td>
<td>79.34%</td>
<td>0.16%</td>
<td>94.03%</td>
</tr>
<tr>
<td>2008</td>
<td>79.07%</td>
<td>0.18%</td>
<td>93.58%</td>
</tr>
<tr>
<td>2006</td>
<td>81.63%</td>
<td>0.23%</td>
<td>91.84%</td>
</tr>
<tr>
<td>2005</td>
<td>81.86%</td>
<td>0.25%</td>
<td>91.51%</td>
</tr>
<tr>
<td>2004</td>
<td>82.97%</td>
<td>0.26%</td>
<td>91.16%</td>
</tr>
<tr>
<td>2003</td>
<td>81.19%</td>
<td>0.30%</td>
<td>90.00%</td>
</tr>
</tbody>
</table>

Notes: Households with a birth include all households where *any* woman aged 15-50 answered yes to the fertility question. Otherwise, the household is classified as not having a birth. Tabulations include households only if the youngest householder’s age is not imputed. All households in US are used in tabulations. Tabulations are unweighted. Source of questions: Q.24 (2011 Survey Instrument) (asked of females aged 15-50); Similar question on other surveys.
<table>
<thead>
<tr>
<th>Table 3</th>
<th>Baby Not Present In Household (Among Women Indicating Birth In Past Year)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Age 20-24</td>
<td>-0.0074 (0.0068)</td>
</tr>
<tr>
<td>Age 25-29</td>
<td>-0.0052 (0.0088)</td>
</tr>
<tr>
<td>Age 30-34</td>
<td>0.0119 (0.0089)</td>
</tr>
<tr>
<td>Age 35-39</td>
<td>0.0566 (0.0113)</td>
</tr>
<tr>
<td>Age 40-44</td>
<td>0.2376 (0.0156)</td>
</tr>
<tr>
<td>Married</td>
<td>-0.0886 (0.0027)</td>
</tr>
<tr>
<td>Income 150-250% FPL</td>
<td>0.0367 (0.0025)</td>
</tr>
<tr>
<td>Income 250-300% FPL</td>
<td>0.0425 (0.0031)</td>
</tr>
<tr>
<td>Income 300%+ FPL</td>
<td>0.0565 (0.0031)</td>
</tr>
<tr>
<td>White</td>
<td>-0.081 (0.004)</td>
</tr>
<tr>
<td>HS Dropout</td>
<td>0.1195 (0.0039)</td>
</tr>
<tr>
<td>HS Graduate</td>
<td>0.0718 (0.0026)</td>
</tr>
<tr>
<td>Non-mover</td>
<td>0.0018 (0.0024)</td>
</tr>
<tr>
<td>Military service</td>
<td>-0.0104 (0.0053)</td>
</tr>
<tr>
<td>Non-citizen</td>
<td>0.0056 (0.0039)</td>
</tr>
<tr>
<td>R²</td>
<td>0.0524</td>
</tr>
</tbody>
</table>

Notes: Sample is based on 242,006 women aged 15-44 giving birth in past year in the US, and is limited to households where there is exactly one woman who indicated giving birth in past year. Baby not present is defined as not having a 0-year-old in household. Households excluded if the youngest member’s age was imputed. In addition to the variables shown above, specifications include state fixed effects and year fixed effects. Omitted categories include Age 15-19, Unmarried, Income 0-150% FPL, Non-white, College Graduate, Mover, Non-military and Citizen.
<table>
<thead>
<tr>
<th></th>
<th>MASS*POST</th>
<th>MASS</th>
<th>POST</th>
<th>N</th>
<th>R²</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>0.0002 (0.0028)</td>
<td>-0.0004 (0.0036)</td>
<td>0.0020 (0.0052)</td>
<td>0.0004 (0.0045)</td>
<td>0.0008 (0.0091)</td>
</tr>
<tr>
<td></td>
<td>-0.0015 (0.0023)</td>
<td>-0.0015 (0.0031)</td>
<td>-0.0123 (0.0039)</td>
<td>-0.0045 (0.0028)</td>
<td>0.0023 (0.0075)</td>
</tr>
<tr>
<td></td>
<td>-0.0014 (0.0021)</td>
<td>0.0000 (0.0026)</td>
<td>-0.0081 (0.0047)</td>
<td>0.0001 (0.0042)</td>
<td>0.0059 (0.0088)</td>
</tr>
<tr>
<td></td>
<td>510,707</td>
<td>79,307</td>
<td>211,135</td>
<td>114,650</td>
<td>96,485</td>
</tr>
<tr>
<td>Fertility rate (pre-reform)</td>
<td>0.0795</td>
<td>0.0146</td>
<td>0.1305</td>
<td>0.0599</td>
<td>0.2011</td>
</tr>
<tr>
<td>Sample</td>
<td>All</td>
<td>Ages 15-19</td>
<td>Ages 20-34</td>
<td>Ages 20-34, Unmarried</td>
<td>Ages 20-34, Married</td>
</tr>
</tbody>
</table>

Notes: All standard errors are clustered at the STATE*YEAR level. The “pre” period is 2003-2006 and the “post” period is 2008-2011. The treatment state is Massachusetts, and the control states are Maine, New Hampshire, Vermont, Rhode Island and Connecticut. Dependent variable is: “Has this person given birth to any children in the past 12 months?” Individual controls included in regression are: education (dropout, HS graduate, college graduate), non-mover, military service, and non-citizen. Women are included in the analysis if they are aged 15-44, resided in New England, and do not have imputed values for gender, fertility, age, marital status, or race.
Table 5
Impact of insurance gains from Massachusetts’ law on fertility

<table>
<thead>
<tr>
<th>INSURED_{S,T,DEMOG}</th>
<th>0.0103 (0.0211)</th>
<th>0.0040 (0.0481)</th>
<th>0.0130 (0.0273)</th>
<th>-0.0459** (0.0216)</th>
<th>0.0982* (0.0504)</th>
<th>-0.0103 (0.0352)</th>
</tr>
</thead>
<tbody>
<tr>
<td>N</td>
<td>510,707</td>
<td>79,307</td>
<td>211,135</td>
<td>114,650</td>
<td>96,485</td>
<td>114,533</td>
</tr>
<tr>
<td>R²</td>
<td>0.0778</td>
<td>0.0394</td>
<td>0.0653</td>
<td>0.0526</td>
<td>0.0155</td>
<td>0.0354</td>
</tr>
<tr>
<td>Fertility rate (pre-reform)</td>
<td>0.0795</td>
<td>0.0146</td>
<td>0.1305</td>
<td>0.0599</td>
<td>0.2011</td>
<td>0.0239</td>
</tr>
<tr>
<td>Sample</td>
<td>All</td>
<td>Ages 15-19</td>
<td>Ages 20-34</td>
<td>Ages 20-34, Unmarried</td>
<td>Ages 20-34, Married</td>
<td>Ages 35-44</td>
</tr>
</tbody>
</table>

Notes: All standard errors are clustered at the STATE*YEAR*DEMOG level. The “pre” period is 2003-2006 and the “post” period is 2008-2011. The treatment state is Massachusetts, and the control states are Maine, New Hampshire, Vermont, Rhode Island and Connecticut. Dependent variable is: “Has this person given birth to any children in the past 12 months?” Individual controls included in regression are: education (dropout, HS graduate, college graduate), non-mover, military service, and non-citizen. Women are included in the analysis if they are aged 15-44, resided in New England, and do not have imputed values for gender, fertility, age, marital status, or race. All specifications include STATE fixed effects (6 categories), YEAR fixed effects (8 categories) and DEMOG fixed effects (48 categories – 2 groups for marital status x 4 groups for poverty status x 6 groups for age status).

** = significant at 5% level, * = significant at 10% level
# Table 6: Sensitivity to set of control variables
Impact of insurance gains from Massachusetts’ law on fertility

<table>
<thead>
<tr>
<th>$INSURED_{ST,DEMOG}$</th>
<th>-0.0459** (0.0216)</th>
<th>-0.0637** (0.0266)</th>
<th>-0.0589 (0.0388)</th>
<th>-0.0492 (0.0434)</th>
<th>0.0982* (0.0504)</th>
<th>0.0998** (0.0501)</th>
<th>0.1144** (0.0544)</th>
<th>0.0988 (0.0664)</th>
</tr>
</thead>
<tbody>
<tr>
<td>N</td>
<td>114,650</td>
<td>114,650</td>
<td>114,650</td>
<td>114,650</td>
<td>96485</td>
<td>96485</td>
<td>96485</td>
<td>96485</td>
</tr>
<tr>
<td>R²</td>
<td>0.0526</td>
<td>0.0537</td>
<td>0.0554</td>
<td>0.0580</td>
<td>0.0155</td>
<td>0.0169</td>
<td>0.0189</td>
<td>0.0222</td>
</tr>
<tr>
<td>Fertility rate</td>
<td>0.0599</td>
<td>0.0599</td>
<td>0.0599</td>
<td>0.0599</td>
<td>0.2011</td>
<td>0.2011</td>
<td>0.2011</td>
<td>0.2011</td>
</tr>
<tr>
<td>Sample</td>
<td>Age 20-34 Unmarried</td>
<td>Age 20-34 Unmarried</td>
<td>Age 20-34 Unmarried</td>
<td>Age 20-34 Married</td>
<td>Age 20-34 Married</td>
<td>Age 20-34 Married</td>
<td>Age 20-34 Married</td>
<td>Age 20-34 Married</td>
</tr>
<tr>
<td>Interaction Terms</td>
<td>STATE, YEAR, DEMOG (Table 4)</td>
<td>STATE*YEAR, DEMOG</td>
<td>STATE<em>YEAR, DEMOG</em>YEAR</td>
<td>STATE<em>YEAR, DEMOG</em>YEAR, DEMOG*STATE</td>
<td>STATE, YEAR, DEMOG (Table 4)</td>
<td>STATE*YEAR, DEMOG</td>
<td>STATE<em>YEAR, DEMOG</em>YEAR</td>
<td>STATE<em>YEAR, DEMOG</em>YEAR, DEMOG*STATE</td>
</tr>
</tbody>
</table>

Notes: All standard errors are clustered at the STATE*YEAR*DEMOG level. The “pre” period is 2003-2006 and the “post” period is 2008-2011. The treatment state is Massachusetts, and the control states are Maine, New Hampshire, Vermont, Rhode Island and Connecticut. Dependent variable is: “Has this person given birth to any children in the past 12 months?” Individual controls included in regression are: education (dropout, HS graduate, college graduate), non-mover, military service, and non-citizen. Women are included in the analysis if they are aged 20-34, resided in New England, and do not have imputed values for gender, fertility, age, marital status, or race.

** = significant at 5% level, * = significant at 10% level
Table 7
Interaction of insurance gains with latent fertility - Impact on fertility

<table>
<thead>
<tr>
<th>Interaction Terms</th>
<th>State</th>
<th>State*Year</th>
<th>State<em>Year</em>Demog2</th>
<th>State<em>Year</em>Demog2*State</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \text{INSURED}<em>{S,T,DEMOG} * \text{FERT}</em>{DEMOG2} )</td>
<td>0.8592** (0.3413)</td>
<td>0.8634*** (0.3419)</td>
<td>0.7980** (0.3838)</td>
<td>0.4198 (0.4638)</td>
</tr>
<tr>
<td>( \text{INSURED}_{S,T,DEMOG} )</td>
<td>-0.0723*** (0.0272)</td>
<td>-0.0742*** (0.0285)</td>
<td>-0.0564** (0.0334)</td>
<td>-0.0297 (0.0382)</td>
</tr>
</tbody>
</table>

\( N \) | 510707 | 510707 | 510707 | 510707 |
\( R^2 \) | | | | |
Mean of dep. variable | All | All | All | All |

Notes: All standard errors are clustered at the STATE*YEAR*DEMOG2 level. The “pre” period is 2003-2006 and the “post” period is 2008-2011. The treatment state is Massachusetts, and the control states are Maine, New Hampshire, Vermont, Rhode Island and Connecticut. Dependent variable is: “Has this person given birth to any children in the past 12 months?” Individual controls included in regression are: education (dropout, HS graduate, college graduate), non-mover, military service, and non-citizen. Women are included in the analysis if they are aged 15-44, resided in New England, and do not have imputed values for gender, fertility, age, marital status, or race. All specifications include STATE fixed effects (6 categories), YEAR fixed effects (8 categories) and DEMOG fixed effects (96 categories – 2 groups for marital status x 4 groups for poverty status x 6 groups for age status x 2 races)

*** = significant at 1% level, ** = significant at 5% level, * = significant at 10% level
Figure 1
Expected Fertility Effects by Age, Marital Status and Gains In Insurance Coverage

Gains in insurance (Income)

<table>
<thead>
<tr>
<th></th>
<th>Big</th>
<th>Small</th>
</tr>
</thead>
<tbody>
<tr>
<td>High Fertility (Age)</td>
<td>↑</td>
<td>no effect</td>
</tr>
<tr>
<td>Low Fertility (Age)</td>
<td>no effect</td>
<td>no effect</td>
</tr>
</tbody>
</table>

High Child Wantedness (Married)

Low Child Wantedness (Unmarried)
Figure 2
Insurance Coverage Rates By Age Group

2a: Massachusetts vs. rest of New England, 2003-2006

2b: Massachusetts vs. rest of New England, 2008-2011

2c: Changes in Coverage Rates in Massachusetts vs. rest of New England, Post-Period vs. Pre-Period
Figure 3
Insurance Coverage Rates By Marital Status
3a: Massachusetts vs. rest of New England, 2003-2006
☐ Rest of New England  ■ Massachusetts

Unmarried  Married

3b: Massachusetts vs. rest of New England, 2008-2011
☐ Rest of New England  ■ Massachusetts

Unmarried  Married

3c: Changes in Coverage Rates in Massachusetts vs. rest of New England, Post-Period vs. Pre-Period
☐ Rest of New England  ■ Massachusetts

Unmarried  Married
Figure 4

Insurance Coverage Rates By Income Group

4a: Massachusetts vs. rest of New England, 2003-2006

4b: Massachusetts vs. rest of New England, 2008-2011

4c: Changes in Coverage Rates in Massachusetts vs. rest of New England, Post-Period vs. Pre-Period
Figure 5
Fertility Rates

5a: Fertility Rates By Age, 2003

5b: Fertility Rates by Age/Marital Status, 2003