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

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Abstract: Health insurance reform in Massachusetts lowered the financial cost of both pregnancy (by increased coverage of pregnancy-related medical events) and pregnancy prevention (by increasing access to reliable contraception and family planning). We examine fertility responses for women of childbearing age in Massachusetts and, on net, find no effect from increasing health insurance coverage. This finding, however, masks substantial heterogeneity. For married women aged 20 to 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 10 percent. Fertility rates changed very little for other groups, in part because of low latent fertility or minimal gains in insurance coverage.

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I. Introduction

Although the Patient Protection and Affordable Care Act (ACA) of 2010 is the first successful attempt in the U.S. to provide near-universal health insurance coverage at the national level, similar policies have been implemented at state and local levels in prior years.[1] Among these regional reforms, the Massachusetts health care law of 2006 – which includes an individual mandate requiring all state residents to obtain health insurance – is the most prominent. Over the years, researchers have used the Massachusetts experience to determine how its new health care law affected health coverage, outcomes, costs, and other critical issues. Given that the Massachusetts legislation served as a model for the design of the ACA, the answers to these questions have broader implications at the national level.

It has been shown that coverage rates increased and out-of-pocket costs of expensive medical events (like pregnancy) decreased as a result of the Massachusetts reform (Long, 2008; Long Stockley, and Yemane, 2009; Kolstad and Kowalski, 2012a).[2] Thus, the reduced cost of pregnancy may have incentivized women of childbearing age who were previously uninsured to plan and carry out a pregnancy. High-cost, anticipated medical events like pregnancy may have been even further subsidized due to the opportunity for adverse selection embedded in the structure of the regulations: even with the law's individual mandate, consumers are given the option to remain uninsured by paying a penalty or purchasing less comprehensive coverage. They can also fairly easily move in and out of more generous plans. Thus, women might

[1] Maine, Vermont, and San Francisco enacted reforms in 2003, 2006, and 2007 respectively. Massachusetts enacted the "Act Providing Access to Affordable, Quality, Accountable Health Care" in 2006. The first three reforms relied on subsidies for purchasing health insurance, while the Massachusetts law also included more far-reaching provisions.

[2] Birth-related expenses in the U.S. ranged from $2,227 to $23,111 in 2010 (Childbirth Connection, http://transform.childbirthconnection.org/resources/datacenter/chargeschart/, Accessed 1/3/2016).

purchase more comprehensive coverage when anticipating pregnancy relative to those that do not plan to have children in the near future.[3] In addition to lowering the out-of-pocket costs of having a baby, the Massachusetts law also lowered the costs of preventing a pregnancy by increasing access to reliable contraception and family planning services. As a result, women who did not want to get pregnant might have increased their use of reliable birth control and thus decreased their fertility rates.

In this paper we use the exogenous changes generated by Massachusetts' health care reform to identify the effect of insurance coverage on fertility. We rely on the American Community Survey (ACS), which explicitly asks questions on fertility. Straightforward difference-in-differences estimates reveal no substantive change in fertility. Since baseline insurance coverage rates varied based on socioeconomic characteristics (rather than just by state and year), we further parameterize the changes in insurance coverage. Even with this parameterized specification, we do not find an effect on realized fertility when we examine all women or stratify the sample by age alone. Our key finding emerges when we stratify by both age and marital status: insurance coverage increased pregnancy for married women aged 20 to 34 by roughly 1 percent and decreased pregnancy for unmarried women of the same age by 10 percent. These opposite-signed results are consistent with different degrees of pregnancy wantedness and different behavioral responses to insurance coverage. These effects cancel out in the aggregate. Fertility for teenagers and older women did not change, which is unsurprising

[3] Feldstein (2013) argues such a design will encourage those who are healthy to strategically remain uninsured until they have a potentially costly medical diagnosis. For example, catastrophic health insurance plans offered under the ACA are available to those under age 30. It covers essential health benefits (including maternity and newborn care) but the high deductibles (approximately $6,300 for an individual) would likely deter many women who anticipate pregnancy from purchasing such a policy. Marton and Yelowitz (2015) find evidence of conditional coverage in the Medicaid system.

since teenagers experienced small gains in insurance coverage (hence, identification is more difficult) and older women have low fertility rates (hence, there is a heterogeneous behavioral response). The results are fairly robust to the inclusion of different sets of control variables and a variety of specification checks.

We also examine and confirm some of the underlying assumptions regarding pregnancy wantedness, physician access, and contraceptive use – all necessary conditions for finding opposite-signed fertility effects. Unlike the fertility results, where we exploit the quasi-experiment of the Massachusetts reform, our analysis cannot establish causality and is thus inherently more speculative due to limited data. Nonetheless, the correlations are consistent with our explanation of fertility patterns, and the overall magnitudes mirror the fertility findings.

The remainder of the paper is arranged as follows: Section II surveys the existing literature on the Massachusetts health insurance reform and fertility responses to expanding health insurance coverage. Section III provides a description of the legislative changes in Massachusetts. Section IV discusses the expected fertility effects from expanding health insurance coverage and shows how the response should vary with observable characteristics. Section V describes the data. Section VI presents the empirical framework and the findings. Underlying assumptions are examined in Section VII, and Section VIII concludes.

II. Literature Review

This paper contributes to an emerging literature evaluating the Massachusetts health care reform, in which insurance coverage and health care utilization are two principal outcomes.[4] Our

[4] Other outcomes include health (Courtemanche and Zapata, 2014; Miller 2012b; Yelowitz and Cannon; 2010), insurance crowd-out (Long, 2008; Miller 2012b; Kolstad and Kowalski, 2012a;

study is the first to examine fertility behavior in this setting. Moreover, unlike existing studies evaluating the fertility effect of insurance mandates, we recognize that marital status (which is broadly consistent with pregnancy wantedness) may differentially affect individuals' responses to newly found health insurance and test this hypothesis in the context of the Massachusetts health reform.

Several studies examine the effect of the Massachusetts reform on insurance coverage (Long, 2008; Long Stockley, and Yemane, 2009; Kolstad and Kowalski, 2012a). There is consensus that coverage rates increased, although there is disagreement on the magnitude.[5] The gains in health insurance coverage varied with socioeconomic characteristics because of heterogeneous baseline coverage: effects were large among young and low-income adults while modest for older and wealthier individuals (Niedzwiecki, 2013). The reform caused little change in coverage for children and teenagers because they were already overwhelmingly eligible under a parent's plan or through Medicaid (Long, Stockley, and Yemane, 2009; Miller, 2012b). The reform also affected healthcare utilization and increased efficiency: the use of preventive health care services increased (Kolstad and Kowalski, 2012a; Niedzwiecki, 2013) and the use of emergency rooms fell (Miller 2012a; Kolstad and Kowalski, 2012a).[6]

Most work focusing on fertility-related moral hazard effects examines Medicaid expansions from the 1980s and 1990s and largely finds a heterogeneous response based on

Yelowitz and Cannon, 2010), labor markets (Kolstad and Kowalski, 2012b), and adverse selection (Hackmann, Kolstrad, and Kowalski, 2012).

[5] Official estimates for the uninsured rate in Massachusetts 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 status if they are uninsured.

[6] Niedzwiecki (2013) finds an overall increase in emergency room visits, however.

demographics.[7] Several studies find different responses by white women (Joyce, Kaestner, and Kwan, 1998; Zavodny and Bitler, 2010; Yelowitz, 1994; DeLeire, Lopoo, and Simon, 2011) and typically no population-wide effect (Zavodny and Bitler, 2010; DeLeire, Lopoo and Simon, 2011). Some also find racial differences in terms of abortion rates (Zavodny and Bitler, 2010; Joyce, Kaestner, and Kwan, 1998; Joyce and Kaestner, 1996). Insurance coverage mandates have also been found to increase the utilization and outcomes of infertility treatments but these results are restricted to older women (Schmidt, 2007; Bitler and Schmidt, 2012).

The increased availability of health insurance also lowers the individual's cost of preventing pregnancy, because almost all health plans cover contraception (and some plans cover abortion). The publicly subsidized "Commonwealth Care" plan in Massachusetts covers a full range of family planning services, including abortion care. Dennis et al. (2012) found that, after the reform, access to affordable contraception improved for low-income women even though they faced new challenges in navigating the system. By providing particular subgroups with a source of entry into the formal health care system, family planning community centers helped overcome such navigation obstacles (Dennis, et al., 2009; Gold, 2009).

Noting that Medicaid has covered contraception since 1972, Kearney and Levine (2009) examined the impact of Medicaid eligibility for family planning services on birth rates and contraceptive use among different demographic groups. They find the largest effects among 20- to 24-year-old women, where birth rates declined by up to 5.1 percentage points (almost 15%) due to higher contraceptive use. Moreover, many studies have shown significant disparities in unintended pregnancies not only by age but also by marital status, education, and income (Finer and Henshaw, 2006; Finer and Zolna, 2011). This literature provides motivation for analyzing

[7] Leibowitz (1990) finds temporary increases in pregnancy rates and births using the RAND health insurance experiment from the 1970s.

fertility responses separately by demographic group because latent fertility (the propensity of a woman to give birth) and pregnancy wantedness vary by sociodemographic characteristics.

In summary, the literature indicates that the Massachusetts reform expanded insurance coverage although the effects differed by socioeconomic groups. The reform also led to increased utilization of health care services. Studies focusing on Medicaid expansions generally find that fertility behavior varies by demographic group, with little evidence for an overall effect. Expanding family planning services, which increases the availability and use of contraception, has been found to be effective in reducing unintended births.

III. Timeline of the Massachusetts Health Care Reform

The Massachusetts health care law dramatically changed the landscape of the state's 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; Kaiser Family Foundation, 2012). 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 of the Federal Poverty Line (FPL). The law also raised enrollment caps for adults and allowed adults younger than 26 years to remain on their parent's plan. In addition, the state provided full coverage to individuals with family incomes up to 150 percent of FPL and subsidized coverage on a sliding-scale basis for those with incomes between 151 and 300 percent of FPL.

The individual mandate, which became effective July 2007, required individuals to purchase health insurance or pay a fine. The penalty was equal to their personal state income tax exemption in the first year and up to 50 percent of the lowest health insurance premium for which they would be eligible in subsequent years. Penalties for not complying with the mandate started in December 2007. The penalty increased from a maximum of $912 per year in 2008 to $1,212 per year by 2011. The penalty only applied to adults, and varied by income and age group; for married couples, the penalty was essentially marriage-neutral, as it was equal to the sum of the individual penalties for each spouse.[8]

For insurance companies, the law stipulated minimum coverage, modified community-rated premiums, and maximum premiums irrespective of preexisting health conditions and claims history. Employers with 11 or more full-time employees were required to offer health insurance or face modest penalties.

IV. Predicted Effects of Expanding Health Insurance Coverage on Fertility

Predicting the effects of health insurance reform on pregnancy is complicated because several other elements, such as 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 family composition and more likely to suffer from infertility.[9] All else equal, single women are less likely to become pregnant because pregnancies are more likely to be unplanned and unwanted. We certainly do not mean to suggest that all pregnancies to married women are wanted and all pregnancies to unmarried women are unwanted. Yet our exploration of this assumption later shows large differences in the reported wantedness of pregnancies. These considerations suggest that if the expansion of health insurance coverage has effects on fertility,

[8] See http://www.mass.gov/dor/individuals/taxpayer-help-and-resources/health-care-reform-information/health-care-reform-regulations-and-tirs/ (Accessed 1/3/2016) for individual mandate penalties for the years 2008 onward.

[9] Around one-third of couples in which the woman is over 35 years old have fertility problems. See http://www.womenshealth.gov/publications/our-publications/fact-sheet/infertility.html (Accessed 1/3/2016).

they would vary by age and marital status. Our hypothesis is that, all else equal, expanding insurance coverage will decrease the fertility of single women (unwanted pregnancies) due to better access to reliable contraception, while increasing the fertility of married women (wanted pregnancies) due to lowering the out-of-pocket cost of pregnancy. In addition, as health insurance becomes more widely available, births by younger women should increase more than those by older women due to the former group's higher latent fertility rates.

The pre-reform baseline distribution of insurance coverage also varied by age and marital status. There were clearly gains in Massachusetts (relative to the rest of New England) from 2008 onward (relative to the period between 2003 and 2006). Moreover, the gains in Massachusetts after the reform varied by family income: some groups experienced minimal gains in insurance (such as relatively affluent women over 300 percent of FPL who were often covered by private insurance), while others experienced much larger gains (such as "near-poor" women with incomes between 150 and 300 percent of FPL).

Larger gains in coverage should lead to larger fertility responses within each age-marital status cell. Figure 1 summarizes the directional expected fertility rate by age, marital status, and income, which roughly reflect variations in latent fertility, child wantedness, and insurance gains. A woman who is young, single (married), and near-poor, would experience larger relative gains in insurance coverage and would be relatively less (more) likely to have a baby after the reform than her more affluent counterpart.

It is also important to highlight 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 teenagers, one might expect smaller fertility responses as well because the insurance gains were typically much smaller.

V. Data Description: American Community Survey, Current Population Survey, and Vital Statistics

Our primary data source is the Census Bureau's ACS Public Use Microdata Sample (PUMS). We use the one-year sample of the ACS PUMS for the years 2003-2011. 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. Moreover, we are able to examine responses for narrow demographic groups, such as married women aged 20 to 34, for which we can more accurately characterize the wantedness of pregnancies and latent fertility. Unlike most household surveys, respondents are required by law to participate in the ACS.[10]

Relevant for our purposes, the ACS directly asks fertility questions for each woman of childbearing age. Specifically, the survey asks, "Has this person given birth to any children in the past 12 months?" Other datasets do not directly ask about fertility; instead, one might impute fertility from the presence of an infant on the household roster. Such an imputation strategy would encounter difficulty in assigning a given infant to a given mother if there was more than one woman of childbearing age in the household. Perhaps more importantly, the ACS reveals that many infants are not living with their mothers: as Table 2 shows, only 81 percent of

[10] Source: https://web.archive.org/web/20121020035539/http://www.census.gov/acs/www/Downloads/language_brochures/ACSQandA_ENG10.pdf (Accessed 1/3/2016). See Title 13, United States Code, Sections 141, 193, and 221. The decennial Census is a notable exception in that it is mandatory.

households where a birth was reported had an infant present. While this non-presence can in large part be attributed to socioeconomic circumstances, some of it simply reflects confusion about the wording of the survey question because the fraction of households reporting a birth who also have a zero- or one-year-old present is 88 percent.[11] Nonetheless, an important difference exists between births and the presence of very young children: roughly 8 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 related to socioeconomic circumstances. We examine 242,006 women aged 15 to 44 who reported a birth (and where that woman was the only one in the household to so report) in the 2003-2011 ACS across the entire U.S. The outcome of interest is whether an infant (defined as age zero) 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, such as grandparents, typically take care of the child. More surprisingly, the likelihood of missing infants increases sharply with age: 35- to 39-year-olds are 6 percentage points more likely to not have an infant present, while 40- to 44-year-olds are nearly 24 percentage points more likely to not have an infant present. These age results should be interpreted differently than the socioeconomic results, however. Fertility is quite low among these age groups – especially 40- to 44-year-olds – and many of the affirmative responses to the fertility question could be related to infant mortality, miscarriage, stillbirth (more likely for older women) or to reporting errors. Given this possibility, we break out our empirical analysis by age group.

[11] 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.

The ACS has one unfortunate drawback: it did not start asking questions on health insurance until 2008, which is the beginning of the "post" period. We rely, instead, on the Current Population Survey (CPS) to derive insurance rates, and append these rates to each woman in the ACS sample. Using the 2004-2012 CPS March Supplements, which cover calendar years 2003-2011, we compute coverage rates for women by demographic category, region, and time.[12] 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 six age groups (15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 39, 40 to 44), four income groups (<150% FPL, 150-250% FPL, 250-300% FPL, >300% FPL), and two marital statuses (married and unmarried). For each demographic group, we create coverage rates for two regions (Massachusetts and the rest of New England) and two periods (the "before" period including calendar years 2003-2006 and the "after" period including calendar years 2008-2011).[13] The total number of groups is therefore 192 (6 ages x 4 incomes x 2 marital statuses x 2 regions x 2 periods). A woman is defined as "uninsured" if she is not covered by private health insurance, Medicare, Medicaid, or

[12] 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. 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 based on age, race, marital status, and education.

[13] 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 the 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 – midway through which the individual mandate was implemented – as a transition year. We exclude the transition year of 2007 when the reform was being phased in, because our interest lies in the effects of the fully phased-in reform; thus we focus on 2008 onward as the "after" period.

CHAMPUS/Tricare military health insurance. The insurance coverage rate is then the 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 teenagers (15- to 19-year-olds) and older women (aged 35 to 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 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- to 24- and 25- to 29-year-olds because young adults leaving college were often no longer covered on a parent's plan and less likely to have a job that provides health insurance coverage. The gains in insurance coverage in Massachusetts following the reform, therefore, were most pronounced for these age groups (Figures 2b and 2c); coverage increased by almost 13 percentage points for 20- to 24-year-olds and 8 percentage points for 25- to 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.

Figure 3 shows that coverage rates are higher for married women than unmarried women, because of the availability of spousal health insurance coverage (Bernstein, et al., 2008). Massachusetts' reform had an equalizing effect for unmarried women: insurance coverage increased by almost 7 percentage points. Figure 4 illustrates the changes in coverage rates by income. Insurance coverage was initially highest for women with incomes over 300 percent of FPL, and the coverage gains were very small (2 percentage points). The coverage gains were also somewhat limited for the poorest women (with incomes less than 150 percent of FPL) because many had health insurance through Medicaid (Sommers, et al., 2012). In contrast, the middle group (with incomes between 150 to 300 percent of FPL) saw increases in insurance coverage of 12 to 16 percentage points.

Finally, although women between 15 and 44 are often categorized as being of childbearing 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 younger women would demonstrate. We calculate a latent fertility variable that represents the propensity of a woman to give birth that varies by age and marital status. To construct the latent fertility rate, we combine two datasets: the 2003 Center for Disease Control and Prevention's (CDC) Vital Statistics data, which records all births in the United States, and ACS data from 2003.

For the numerator, we use CDC's 2003 natality data to establish baseline fertility rates for 12 demographic cells – six age groups (15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 39, and 40 to 44) and two marital statuses (married and unmarried). For the denominator, we use ACS data from 2003 to obtain the total number of women within each demographic cell.[14]

To compute the latent fertility variable for married women aged 20 to 24 years, for example, we divide the number of births from women in this demographic cell (483,843) by the total number of women in the U.S. within this same cell (2,255,895) and obtain a latent fertility rate of 21.4 percent. The inverted-U shape of latent fertility in Figure 5a illustrates wide variations in the propensity for having a baby, with women aged 20 to 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). Although not shown, conditional on age and marital status, race

[14] Results were virtually identical using data from 2011.

is not an important factor affecting latent fertility. These fertility rates provide strong motivation for stratifying the sample, both by age alone and by age and marital status.

VI. Empirical Framework and Results

VI.1 Empirical Framework

As is well recognized, the Massachusetts reform creates a quasi-experiment to evaluate the impact of expanding health insurance coverage. The natural starting point for our examination of fertility is a straightforward difference-in-differences (DD) estimator estimated from a linear probability model:[15]

$$BIRTH_{ijt} = \beta_0 + \beta_1 MASS_j * POST_t + \beta_2 MASS_j + \beta_3 POST_t + \beta_4 X_{ijt} + \beta_5 UR_{ijt} + \varepsilon_{ijt} \quad (1)$$

where $BIRTH_{ijt}$ is an 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). The vector X_{ijt} includes controls for the woman's education (high school dropout, high school graduate, college graduate is omitted), whether the woman has changed residence in the past year, whether she has served in the military, race/ethnicity (African-American/Black, Hispanic/Latino, Other non-white), and whether she is a non-U.S. citizen. The specification also controls for the age groups, income groups, and marital status. Finally, UR_{ijt} is the age-specific unemployment rate for women in each state-year cell, created from the ACS; transitory changes

[15] Results from a probit model are similar.

in wages can affect the timing of fertility (Dehejia and Lleras-Muney, 2004). The coefficient estimate on β_1 is then interpreted as the DD estimator.[16]

Although transparent, there are reasons to go beyond the specification in equation (1). Most importantly, although the near-universal health reform in Massachusetts leveled coverage rates across groups, there were very different gains based on a woman'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.[17] Thus, equation (2), which forms our baseline specification of insurance gains on fertility, is:

$$BIRTH_{idjt} = \beta_0 + \beta_1 INSURED_{djt} + \beta_2 DEMOG_d + \beta_3 X_{ijt} + \delta_s + \delta_t + \varepsilon_{ijt} \tag{2}$$

where $INSURED_{djt}$ is the fraction of demographic group *d* covered in region *j* in period *t*.[18] It is likely that the key components of $INSURED_{djt}$ – especially demographics such as 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 (δ_s and δ_t). The estimate of the impact of insurance coverage, β_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 sets of controls.

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[16] The large majority of papers studying the effect of the Massachusetts healthcare law use some form of difference-in-differences identification strategy. See, for example, Kolstad and Kowalski (2012a), Courtemanche and Zapata (2014), Yelowitz and Cannon (2010), Long, Stockley, and Yemane (2009), and Miller (2012a).

[17] 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, Lopoo, and Simon, 2011; Zavodny and Bitler, 2010; Currie and Gruber, 2001).

[18] All standard errors are clustered at the STATE level.

One key drawback to equation (2) 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. Older women 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 or aborting a pregnancy. One would expect that pregnancies are much more likely to be unwanted for single women, and wanted for married 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.[19]

VI.2 Basic Results

The full sample consists of more than 500,000 women aged 15 to 44 in Massachusetts and surrounding states. Summary statistics are shown in Appendix Table 1 for the full sample and various subgroups. Nearly 8 percent reported a birth in the past year. 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. Roughly 13 percent of women aged 20 to 34 had a baby in the previous 12 months, a much higher rate than for women aged 15 to 19 or 35 to 44. The fertility differences are especially pronounced by marital status; approximately 20 percent of married women aged 20 to 34 reported having a baby, more than three times the rate

[19] Joyce, Kaestner, and Kwan (1998) only include young, single, and low-educated women in their sample. Zavodny and Bitler (2010) stratify their sample by race/marital status and race/education to analyze the fertility effects.

of unmarried women in the same age group. The summary statistics show that prior to reform, the samples in Massachusetts and the remaining states are extremely similar. Pregnancy rates were roughly 8 percent in both areas, insurance coverage was 91 percent, and the unemployment rate was 7 percent. There are very small differences in racial composition, marital status, or the age distribution. Income and education levels in Massachusetts are somewhat higher than other states, but all regression specifications will control for those factors.

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 that the expansions in insurance had little effect on fertility. In all cases, the coefficient estimate is substantively small and, with the exception of 35 to 44 year-old women, statistically 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 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 little effect on overall fertility. Yet, as shown in columns (4) and (5), there are opposite-signed effects for unmarried and married women aged 20 to 34. Although not shown, coefficient estimates are insignificant and much smaller for other age/marital status groups. For unmarried women aged 20 to 34, insurance coverage increased by

11.5 percentage points due to the Massachusetts law.[20] With a coefficient estimate of -0.0531, this would imply that fertility fell by -0.61 percentage points. Since the pre-reform baseline fertility in the ACS was 5.98 percent, then fertility fell by 10.1 percent. This result is similar in magnitude to Kearney and Levine (2009) who find that Medicaid eligibility for family planning services led to a 15 percent decline in birth rates for 20-24 old women. For married women in the same age group, our results indicate that gains in insurance coverage led to increased fertility. The overall gain in insurance coverage was 2.5 percentage points, leading to an increase in fertility of 0.25 percentage points from a much higher baseline of 20.1 percent. Thus, among married women, fertility increased by around 1.2 percent.[21]

The model in equation (2) assumes instant adjustment to newfound health insurance coverage. Although the individual mandate in Massachusetts started on July 1, 2007, the penalties for not complying with the mandate began at year-end and enrollment soared in December 2007 (Chandra, Gruber, and McKnight, 2011). This delayed enrollment, combined with lags in getting pregnant, delivering a baby, or obtaining birth control, suggests that the effects on fertility may be larger in post-reform years after 2008. We estimated the model in Table 5 – excluding 2008 – and find somewhat larger effects. For unmarried women, the coefficient (standard error) is -0.0644 (0.0155), or 21 percent larger. For married women, the estimates are 0.1097 (0.0206), or 11 percent larger. Thus, our baseline estimates appear to understate the longer-run effect on fertility. It should be noted, however, that there could have been pent-up demand for pregnancies that will slow down after a few years. If that is the case,

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[20] We ran difference-in-difference estimates similar to equation (1) to calculate the change in insurance coverage.
[21] It is possible that the positive effects for married women are related to newly-found coverage of infertility treatment. This is not very likely, however, since infertility tends to predominantly affect women over 35 years old.

our baseline estimates would be somewhat larger than the long-run fertility effects of the health reform.

VI.3 Extensions

Given the striking differences in insurance coverage for 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. 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; penalties for married couples who did not comply with the individual mandate rules (with or without children) equaled the sum of the individual penalties for each spouse. To examine this further, we run difference-in-differences regressions in which the outcome is whether the woman is married; for both the full sample as well as each age group, the estimate is insignificant.[22]

One concern about Table 5 is that the INSURED variable is a complicated function that incorporates demographics, state, and year. Although we control for each of the main effects, interactions of DEMOG, STATE, and YEAR may directly impact 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

[22] Our specifications look at changes in the proportion married. It is possible that the Massachusetts mandate affects flows in and out of marriage. Unfortunately, the ACS only asks about getting married and getting divorced/separated/widowed from 2008 onward. Abramowitz (forthcoming) finds effects of the ACA young adult provision on marriage.

those differences – rather than health insurance coverage – could drive fertility decisions. In Table 6, we test the sensitivity of the coefficient estimates to these kinds of concerns for women aged 20 to 34. For comparison, column (1) replicates the specification from the prior table. In columns (2) and (3) we include state-specific time trends and STATE*YEAR effects, and observe that their inclusion have little impact on the underlying conclusions (if anything, the negative impact is stronger for unmarried women). Column (4) adds DEMOG*YEAR interactions (in addition to the STATE*YEAR interactions). The coefficient estimates have a slightly higher magnitude. Finally, column (5) 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). It is reassuring that the actual coefficients are quite similar to the baseline specification, and the married sample retains statistical significance. Overall, the findings for both unmarried women and married women hold up well to including additional controls.

Another related concern about the specification is that differential, pre-existing, pre-program trends in fertility might exist in Massachusetts and other states, and the specifications in Tables 4 and 5 do not control for them. Figure 6 illustrates fertility rates for married and single women, aged 20 to 34, and at first blush, there appear to be no differences prior to the health insurance reform.[23] Nonetheless, we control for the possible presence of such pre-program trends in our regressions by estimating a linear time trend for each of 72 demographic groups (i.e., 6 ages x 2 marital statuses x 6 states; such as 20- to 24-year-old married women in Massachusetts) based only on pre-program fertility (i.e., using the years 2003 to 2006) in the ACS and generate predicted trend values for 2003 to 2011 (Chakrabarti and Roy, 2011). We then additionally

[23] The lack of a fertility effect in the “post” period in Figure 6 is consistent with the findings in Table 4, and illustrates the importance of modelling the gains in insurance coverage.

control for fertility trends in the model. The results for health insurance coverage are qualitatively similar incorporating such trends. For example, the impact of health insurance coverage on unmarried women aged 20 to 34 is negative (and statistically significant), with a coefficient (standard error) of -0.0475 (0.0148). For married women aged 20 to 34, the coefficient (standard error) is 0.0888 (0.0171).

There may also be a concern that the generous health insurance benefits in Massachusetts – with community-rated premiums and guaranteed issue – make the state a more attractive place for individuals with high expected medical costs – such as pregnant women – and, therefore, encourages migration. 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 theoretical concern.[24] 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 to 29. The result is consistent with a greater implicit tax on the young arising from community rating and individual mandates. The same factors that generate implicit taxes for the majority of 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 non-movers. Our results are quite similar to the baseline results in Table 5. For the 94 percent of unmarried women aged 20 to 34 who did not move across state lines, the coefficient estimate (standard error) is now -0.0456 (0.0148) compared with the initial estimate

[24] 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 (2014) account for migration responses due to Medicaid.

of -0.0531 (0.0138). For the 95 percent of married women aged 20 to 34 who did not move across state lines, the coefficient estimate (standard error) is now 0.1014 (0.0285) compared with the initial estimate of 0.0992 (0.0235). Thus, the basic conclusions remain unchanged by restricting the sample to non-movers.

Another concern is that some of the comparison states – in particular Maine and Vermont – made changes to their health care system. Maine enacted the "Dirigo Health Reform" in 2003 and Vermont adopted "Catamount Health" in 2006, both of which subsidized the purchase of health insurance. Although Table 6 shows the fertility results for both single and married women are robust to the inclusion of state-trends or STATE*YEAR effects, we have also re-estimated the specification from Table 5 without Maine and Vermont. For single women aged 20 to 34, the coefficient estimate (standard error) is now -0.0587 (0.0157) and for married women the estimates are 0.0975 (0.0250). Comparing these coefficient estimates to those in column (1) of Table 6, one can see that the exclusion of these states from the control group has little impact, a finding consistent with Kolstad and Kowalski (2012a) and Courtemanche and Zapata (2014).

It should also be noted that the grouping of women between 20 and 34 may be too large, especially since marital status varies considerably within this grouping. Moreover, many women in the 20 to 24 age range may be enrolled in college full-time, and provisions for covering college students could vary by state and year. We therefore ran regressions separately for younger women in this group, as well as "prime-aged" women with respect to fertility (between 25 and 34, inclusive; see Dehejia and Lleras-Muney, 2004). The overall pattern of results is consistent with the findings in Table 5, and provides little justification for breaking out the sample further. For example, for both college-aged women and prime-aged women, the overall effect of insurance coverage is insignificant. For unmarried college-aged women and prime-aged women, the effect of expanding health insurance is to reduce fertility, with respective coefficient estimates (standard errors) of -0.0479 (0.0251) and -0.0807 (0.0268). For married college-aged women and prime-aged women, expanding insurance increases fertility, with respective coefficient estimates (standard errors) of 0.1591 (0.0547) and 0.0633 (0.0368).

One might also argue that the Massachusetts reform should have no effect on the fertility of single women given the existence of the Title X network of family planning clinics which have provided, and continued to provide after the reform, women's health services, including contraception. It should be noted that Title X clinics' clients receive care at minimal or no cost since they are predominantly low-income - 67 percent of them have family incomes less than 100% of FPL and 90 percent have incomes less than 200% of FPL. However, individuals with incomes above 250% of FPL have to pay the full price of care[25, 26]. Given that the largest gains in insurance coverage were experienced by individuals with incomes 250-299% of FPL (Figure 4c), we can hypothesize that the observed negative fertility effect is driven by this subgroup while fertility for lower-income women remained closer to pre-reform levels.

Finally, we have examined an alternative definition of fertility. Recall that some women reported pregnancies but did not have infants living in the household, and on many dimensions the missing infants varied in logical ways with socioeconomic characteristics, such as marital status, race, and education. We have run similar specifications to our baseline result, but where childbirth is now defined as having an infant on the household roster. Such a measure creates difficulty in linking the infant to a mother when there are multiple women of childbearing age in a household, or when the mother is absent from the household. Although we continue to find

[25] See http://www.nationalpartnership.org/research-library/repro/title-x-backgrounder.pdf (accessed 1/3/2016)
[26] See http://www.guttmacher.org/pubs/gpr/16/3/gpr160314.html (accessed 1/3/2016)

reductions in fertility for unmarried women aged 20 to 34 and increases in fertility for married women aged 20 to 34, the coefficient estimates are roughly one-quarter to one-half as large and not statistically significant. This provides further evidence supporting 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 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 to 34, and also explain the non-existence of effects for other groups. Married women in this age bracket increased their fertility when experiencing gains in insurance coverage because pregnancies are largely wanted and underlying fertility is high. Single women, on the other hand, decreased their fertility because pregnancies are largely unwanted and better access to contraception helps them prevent or terminate pregnancy. For women aged 35 and older, latent fertility is relatively low (and insurance coverage was typically high prior to the reform), so the overall fertility responses are small. For teenagers, fertility rates are also quite low, many pregnancies are unwanted, and insurance coverage was fairly high prior to the reform. Thus, we find small effects for them, too.

VII. Exploring the Underlying Assumptions

In this section we investigate two key relationships that lead to different fertility responses: the relationship between marital status and pregnancy wantedness and the relationship between insurance coverage and birth control methods. As we do so, an important caveat should be kept in mind: for neither relationship do we have adequate data to exploit the same quasi-experiment as we did for the fertility results. As such, the results in this section should be thought of as correlations and cannot establish causality.

VII.1 Pregnancy Wantedness

The explanation of our results rests on the assumption that children born to married women are typically wanted, while children born to single women are typically unwanted. We explore this with data from the Pregnancy Risk Assessment Monitoring System (PRAMS), an annual survey conducted by the CDC of women who had a live birth. New mothers participated within four months of giving birth and are asked about their attitudes towards the pregnancy.[27] We obtained data for Massachusetts, Maine, Rhode Island, and Vermont; Connecticut and New Hampshire did not participate in PRAMS. Massachusetts data is only available between 2007 and 2010 (which makes it impossible to estimate difference-in-differences models as we did with fertility), while data for the other states are available from 2003 to 2011.

Respondents are asked to report pregnancy wantedness right before becoming pregnant. Figure 7a shows that among married women who gave birth, 79 percent reported wanting to have a baby; in contrast, the majority of single women reported unwanted pregnancies. It is likely that pregnancy wantedness is overstated for both groups of women because PRAMS does not survey women who had abortions. Forty percent of all unwanted pregnancies end in abortions, and 45 percent of all abortions are obtained by single women.[28] A back-of-the-envelope calculation

[27] There are two parts to the survey: the core questions and state-specific questions. The core questions include attitudes and feelings about the most recent pregnancy, content and source of prenatal care, maternal alcohol and tobacco consumption, physical abuse before and during pregnancy, pregnancy-related morbidity, infant health care, and contraceptive use.
[28] See http://www.guttmacher.org/pubs/fb_induced_abortion.html (Accessed 1/3/2016).

using abortion statistics from the Guttmacher Institute implies that the proportion of unwanted pregnancies for single women is 68 percent (Figure 7b).[29]

We also attempted to examine the changes in abortion rates for married and single women in Massachusetts and the control states over the study period. If our results hold, we would expect to see fewer abortions for young single women in Massachusetts relative to the rest of New England after the reform. Unfortunately, the CDC's annual "Abortion Surveillance" reports provide abortion data by marital status for only two of our control states before 2007 and do not provide any such data for Massachusetts after 2007.

Next, we evaluate the impact of marital status on pregnancy wantedness for live births. Pregnancy wantedness equals "one" if the woman wanted to be pregnant right before becoming pregnant or sooner, and equals "zero" if the woman wanted to become pregnant later or did not want to become pregnant at all (Finer and Kost, 2011). Using a linear probability model, we regress pregnancy wantedness on marital status and incrementally add explanatory variables (mother's age, education, race, number of previous live births, and insurance status) in Table 7.[30]

[29] The PRAMS shows the number of unwanted live births to married and single women to be 82,712 and 133,473, respectively. Since 60 percent of unintended pregnancies result in a birth and 40 percent result in abortion, the total number of unintended pregnancies is 360,308, of which there are 144,123 abortions. Approximately 45 percent of abortions are obtained by single women (64,855 abortions), while 55 percent are obtained by married women (79,268 abortions). Adding abortions to unwanted births reveals that 68 percent of all pregnancies for unmarried women are unwanted (compared with 59 percent when examining live births), while 34 percent of all pregnancies for married women are unwanted (compared with 21 percent when examining live births).

[30] Marital status is equal to 1 if the mother is married and 0 otherwise. There are 7 categories for mother's age: less than 17 years old, 18-19 years old, 20 to 24 years old, 25 to 29 years old, 30 to 34 years old, 35 to 39 years old, and over 40 years old. There are 5 categories for mother's education: 0-8 years, 9-11 years, 12 years, 13-15 years, and 16 years or more. Mother's race is equal to 1 if the mother is white and 0 otherwise. A woman is considered insured right before she became pregnant if she had any type of health insurance including Medicaid. All specifications include state fixed effects.

Pregnancy wantedness is at least 20 percentage points higher for married women and is very stable regardless of specification.

A natural concern with the PRAMS is the issue of reporting bias, and possibly differential reporting bias by marital status. If women perceive reporting pregnancy unwantedness as socially undesirable, they may report unwanted pregnancies as wanted pregnancies. Dietz, et al. (1999) note that the PRAMS consists primarily of self-administered mailed questionnaires, and argue that many women may find the self-administered questionnaire a less threatening forum for reporting an unintended pregnancy. The same study finds unadjusted differences in unintended pregnancies of roughly 30 percentage points by marital status across eight states, using data from roughly 15 years before our PRAMS analysis. The way in which the PRAMS is administered, along with the stability of the marital gap in wantedness across time, states, age groups, and empirical specifications suggests genuine differences in pregnancy wantedness rather than reporting issues.

VII.2 Health Insurance Coverage, Physician Access, and Contraceptive Usage

We next explore underlying behavioral responses to health insurance coverage, in particular whether single women switch to more reliable contraception methods (which typically require a doctor's prescription). To clearly demonstrate the link between health insurance and contraceptive usage we adopt a two-step approach. First, we investigate whether health insurance leads to greater physician access. Second, we analyze the effect of physician access on the use of different contraceptive methods.

For this analysis we use the CDC's Behavioral Risk Factor Surveillance System (BRFSS). The BRFSS is a telephone survey of personal health behaviors. The core questionnaire asks about health insurance and primary care physician access.[31] Family planning questions are available intermittently (for the years 2004, 2006, 2010, and 2011) and only asked for respondents in all states in 2004.[32] We therefore restrict our analysis to the 2004 BRFSS and focus on the states in New England, all of which had a contraception coverage mandate for health insurance plans by 2004.[33]

We first examine the relationship between the availability of health insurance and access to a primary care provider. Health insurance coverage equals "one" if the person has any kind of health care coverage, or "zero" otherwise. Physician access equals "one" if the respondent reported that she thinks of one or more people as a personal doctor or health care provider. All specifications examine women of childbearing age, but exclude women who use permanent contraceptive methods or have partners who use permanent contraceptive methods.[34] Table 8 estimates a linear probability model, where access to a health care provider is regressed on health insurance coverage, while incrementally controlling for additional factors in the remaining columns. The coefficient on health insurance is positive, highly significant, and implies a 20-percentage point increase in access. This effect is present for the full sample, as well as for

[31] The question about health insurance coverage is: "Do you have any kind of healthcare coverage, including health insurance, prepaid plans such as HMOs, or government plans such as Medicare?" The question about doctor access is: "Do you have one person you think of as your personal doctor or healthcare provider?"

[32] The family planning module was included in the survey by the following states: Arizona, Kentucky, Minnesota, Missouri, Montana, Oregon, and Wisconsin in 2006; Delaware, Florida, Kentucky, Mississippi, and Montana in 2010; Arizona, South Carolina, and Tennessee in 2011.

[33] These states mandated contraception coverage between 1999 and 2003 (Mulligan, 2014; Dills and Grecu Cotet, 2014).

[34] The median age of women using permanent contraception is higher than the median age of women using temporary forms of birth control (37 versus 32), since completed fertility is positively correlated with age. Agüero and Marks (2011) discuss measurement error in self-reported measures of fertility, such as the fear of revealing fertility problems to a survey taker. Reporting issues would likely lead to attenuation bias in the regression models on contraception use.

married and unmarried women; insurance coverage leads to a larger gain in access to physicians for single women.

Having established that insurance coverage is positively associated with physician access, we next examine the role of access on contraceptive use. The top half of Table 9 shows – for all women as well as by marital status – no relationship between physician access and overall contraceptive use. Having access to a physician has insignificant – and substantively small – effects on contraceptive use. The bottom half of the table examines the use of reliable contraceptive methods. Contraceptive methods that have less than a 12 percent typical failure rate are classified as effective, and include intrauterine devices, implants, shots, pills, contraceptive patches, diaphragms, and cervical rings and caps.[35] Condoms, emergency contraception, withdrawal, and rhythm are some of the less effective contraceptive methods. Having access to a primary care physician is positively associated with using more effective contraceptive methods, of approximately 13 percentage points. The effect for single women is generally larger than the effect for married women, and the magnitudes suggest substantial overall shift towards more reliable contraceptive methods.

The contraception findings are consistent with the findings on pregnancy wantedness: if a single woman obtains health insurance and better access to physicians, she is more likely to switch to reliable contraception because pregnancy wantedness is lower. As importantly, the

[35] 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 requires a doctor's prescription or contact with a healthcare provider. 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 healthcare providers. See https://web.archive.org/web/20110302033458/http://www.plannedparenthood.org/health-topics/birth-control//birth-control-effectiveness-chart-22710.htm, (Accessed 1/3/2016).

overall magnitudes are consistent with the reduction in fertility for single women. Recall that insurance coverage increased by 11.5 percentage points for single women aged 20 to 34. Such gains would lead to a roughly 2.9 percentage point increase in access to primary care physicians (Table 8). And the increase in access then leads to roughly a 0.44 percentage point increase in reliable contraception use (Table 9). This is the same order of magnitude as the reduction in fertility of -0.61 percentage points (Table 5).

VIII. Conclusions and Discussion

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 to 34, this ignores substantial heterogeneity across married and unmarried women (which proxies for child wantedness). Among young single women, fertility decreased by 10 percent while fertility increased by 1 percent for young married women. We find no effect on birth rates for teens or older women.

Whether the reform shifted the timing of births or changed the total number births remains an open question. Evidence from other policy contexts suggests the importance of timing considerations (Dickert-Conlin and Chandra, 1999; Gans and Leigh, 2009; Neugart and Ohlsson, 2013; LaLumia, Sallee, and Turner, 2013; Schulkind and Shapiro, 2014). Furthermore, abortion and birth control access have been found to affect life-cycle fertility in the U.S. and abroad (Ananat, Gruber, and Levine, 2007; Pop-Eleches, 2010). Data over a longer period are needed to assess the long-run effects. Regardless of whether the reforms reflect timing or level effects, the proportion of unintended pregnancies – those that are mistimed, unplanned, or unwanted – fell as a result of the reform.

Our results have implications for the ACA. Expanding insurance would likely increase wanted pregnancies on a national level and decrease unwanted births. There are three reasons to believe the fertility reductions for single women in Massachusetts were smaller than what would occur from the ACA. First, even prior to reform, health insurance coverage was quite high in Massachusetts; nearly 90 percent of the Massachusetts population had insurance compared with roughly 85 percent for the U.S. population (U.S. Census Bureau, 2007). Larger changes in insurance coverage would lead to larger reductions in fertility for single women. Second, abortion and family planning services are more accessible in Massachusetts even without insurance (Guttmacher Institute, 2015), leading to a smaller role for fertility reductions due to effective contraception. Third, the fraction of pregnancies that are unintended (unwanted or mistimed) is lower in Massachusetts than many other states (Finer and Kost, 2011). The combination of these factors – larger insurance gains, lower access to family planning, and greater unwantedness – suggests a larger impact from the set-up of the health insurance marketplaces in 2014 and the employer mandate in 2015 from the ACA. There are two possible limiting factors. First, in Massachusetts some previously uninsured individuals gained coverage through the Medicaid expansion while with the ACA states can choose not to expand Medicaid.[36] However, our estimates remain largely applicable to the ACA since for 2006-2010 only 24% of the Medicaid enrollment growth in Massachusetts was related to the expansion.[37] Second, although under ACA all new health plans must cover certain women's preventive services with no co-payments, including contraceptive counseling and the full range of FDA

[36] As of December 17, 2015 31 states have adopted the Medicaid expansion. See http://kff.org/health-reform/state-indicator/state-activity-around-expanding-medicaid-under-the-affordable-care-act/, (Accessed 1/3/2016)

[37] See http://bluecrossfoundation.org/sites/default/files/MassHealth%20Enrollment%20Growth%20May%202011.pdf, (Accessed 1/3/2016)

approved contraception methods, the Supreme Court's decision in *Burwell v. Hobby Lobby* exempted closely held corporations from providing coverage of contraception if such provisions violate the owners' religious beliefs. Although a non-trivial share of employment in the private sector derives from closely held corporations, large employers overwhelmingly offered contraception prior to the mandate (Griswold, 2014), suggesting the actual impact would be minor.

A reduction in unwanted births – either in fertility levels or timing and spacing – could have favorable implications for child investment (Buckles and Munnich, 2012) or societal outcomes such as criminal activity (Donohue and Levitt, 2001). Reducing unwanted births may also lead to increased investment in a woman's own human capital and the human capital of her children. Ultimately, the potential savings may be far greater than the financial resources currently spent on unintended pregnancies, at both the level of the individual and society.

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Table 1: Implementation Timeline for the Massachusetts Health Care Reform	
April 2006	• Health Care Reform legislation passed
July 2006	• Federal Government approves Medicaid waiver for health care reform
October 2006	• Plan Type I for Commonwealth Care open for enrollment (for residents at 100% of FPL)
January 2007	• Plan Types II, III and IV for Commonwealth Care open for enrollment (for residents between 100% and 300% of FPL)
March 2007	• Deadline for Connector Board to set minimum "creditable" coverage standards
May 2007	• Commonwealth Choice plans become available (individuals and small businesses can buy insurance)
July 1, 2007	• Individual mandate to purchase health insurance becomes effective • Deadline for employers to provide health insurance to full-time employees • Deadline for individual and small-group insurance markets to merge
January 2008	• Individual mandate penalty becomes effective: 50% of premium per month if uninsured

Sources: See The Henry J. Kaiser Family Foundation, "Focus on Health Reform," June 2007, accessed 1/3/2016 from http://kaiserfamilyfoundation.files.wordpress.com/2013/01/7494-02.pdf ; State of Massachusetts, "Massachusetts Health Care Reform Fact Sheet," accessed 1/3/2016 from https://web.archive.org/web/20150316101109/http://www.mass.gov/eohhs/docs/mrc/health-care-reform-fact-sheet.rtf

Table 2: Comparison of Fertility Reports versus Infants on Household Roster			
	Is there an infant on household roster?		Is a birth reported?
	Conditional on reporting a birth in last 12 months	Conditional on not reporting birth in last 12 months	Conditional on having an infant on household roster
All Years	81.14%	0.21%	92.24%
2011	78.52%	0.18%	92.26%
2010	80.17%	0.18%	92.89%
2009	80.73%	0.17%	93.74%
2008	80.60%	0.19%	93.28%
2006	82.49%	0.24%	91.56%
2005	82.82%	0.26%	91.26%
2004	83.83%	0.27%	90.93%
2003	82.08%	0.30%	89.81%

Notes: Sample drawn from the 2003-2011 ACS. Households reporting a birth include all households where any woman aged 15-44 answered yes to the ACS fertility question: “Has this person given birth to any children in the past 12 months?” 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 the U.S. are used in tabulations. Tabulations are unweighted. Source of questions: Q.24 (2011 Survey Instrument) (asked of women of childbearing age). Similar question on other surveys.

Table 3: Baby Not Present (In Households with Woman Reporting Birth)	
Age 20 to 24	-0.0074 (0.0068)
Age 25 to 29	-0.0052 (0.0088)
Age 30 to 34	0.0119 (0.0089)
Age 35 to 39	0.0566*** (0.0113)
Age 40 to 44	0.2376*** (0.0156)
Married	-0.0886*** (0.0027)
Income 150-250% FPL	0.0367*** (0.0025)
Income 250-300% FPL	0.0425*** (0.0031)
Income 300%+ FPL	0.0565*** (0.0031)
White	-0.081*** (0.004)
High School Dropout	0.1195*** (0.0039)
High School Graduate	0.0718*** (0.0026)
Non-mover	0.0018 (0.0024)
Military service	-0.0104** (0.0053)
Non-citizen	0.0056 (0.0039)
R^2	0.0524

Notes: Sample drawn from the 2003-2011 ACS. Standard errors in parentheses. Sample is based on 242,006 women aged 15-44 giving birth in past year in the U.S., and is limited to households in which exactly one woman indicated she had given birth that year. "Baby not present" refers to a household that does not have a zero-year-old. Households were 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 to 19, Unmarried, Income 0-150% FPL, Non-white, College Graduate, Mover, Non-military and Citizen.

*** = significant at 1% level, ** = significant at 5% level

Table 4: Difference-in-differences Estimates of the Impact of Health Reform on Fertility						
MASS*POST	-0.0000	-0.0009	-0.0016	0.0016	-0.0020	0.0027**
	(0.0026)	(0.0020)	(0.0058)	(0.0034)	(0.0081)	(0.0010)
MASS	0.0032	-0.0009	0.0021	-0.0034*	0.0078	0.0080*
	(0.0035)	(0.0012)	(0.0063)	(0.0015)	(0.0117)	(0.0035)
POST	-0.0009	-0.0003	0.0024	-0.0010	0.0062	-0.0029*
	(0.0027)	(0.0020)	(0.0058)	(0.0038)	(0.0081)	(0.0013)
N	507,000	78,763	209,477	113,701	95,776	218,760
R^2	0.0650	0.0378	0.0590	0.0529	0.0112	0.0322
Fertility rate (pre-reform)	0.0794	0.0147	0.1303	0.0598	0.2012	0.0554
Sample	All	Ages 15 to 19	Ages 20-34	Ages 20-34, Unmarried	Ages 20-34, Married	Ages 35 to 44

Notes: Sample drawn from the 2003-2011 ACS (excluding 2007). Standard errors in parentheses. All standard errors are clustered at the STATE 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, high school graduate, college graduate), non-mover, military service, race/ethnicity, age category, marital status, income category and non-citizen. The unemployment rate – measured by state/year/age group/marital group – is also included. 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.

** = significant at 5% level, * = significant at 10% level

Table 5: Impact of Insurance Gains on Fertility

$INSURED_{djt}$	0.0082	0.0077	0.0093	-0.0531**	0.0992***	-0.0048
	(0.0150)	(0.0226)	(0.0121)	(0.0138)	(0.0235)	(0.0371)
N	507,000	78,763	209,477	113,701	95,776	218,760
R^2	0.0783	0.0400	0.0664	0.0556	0.0175	0.0357
Fertility rate (pre-reform)	0.0794	0.0147	0.1303	0.0598	0.2012	0.0554
Sample	All	Ages 15 to 19	Ages 20-34	Ages 20-34, Unmarried	Ages 20-34, Married	Ages 35 to 44

Notes: Sample drawn from the 2003-2011 ACS (excluding 2007). Standard errors in parentheses. All standard errors are clustered at the STATE 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, high school graduate, college graduate), non-mover, military service, race/ethnicity, and non-citizen. The unemployment rate – measured by state/year/age group/marital group – is also included. 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 1% level, ** = significant at 5% level, * = significant at 10% level

Table 6: Sensitivity Analysis					
	Age 20-34, Unmarried (N=113,701; Pre-reform fertility rate=0.0598)				
$INSURED_{djt}$	-0.0531**	-0.0608**	-0.0788**	-0.0914*	-0.0645
	(0.0138)	(0.0234)	(0.0267)	(0.0357)	(0.0374)
R^2	0.0556	0.0558	0.0568	0.0586	0.0613
	Age 20-34, Married (N=95,776; Pre-reform fertility rate=0.2012)				
$INSURED_{djt}$	0.0992***	0.0967**	0.1004***	0.1113***	0.1018**
	(0.0235)	(0.0252)	(0.0243)	(0.0223)	(0.0272)
R^2	0.0175	0.0176	0.0188	0.0209	0.0242
Interaction Terms	STATE, YEAR, DEMOG (Table 5)	STATE, YEAR, State trends, DEMOG	STATE*YEAR, DEMOG	STATE*YEAR, DEMOG*YEAR	STATE*YEAR, DEMOG*YEAR, DEMOG*STATE

Notes: Sample drawn from the 2003-2011 ACS (excluding 2007). Standard errors in parentheses. All standard errors are clustered at the STATE 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, high school graduate, college graduate), non-mover, military service, race/ethnicity, and non-citizen. The unemployment rate – measured by state/year/age group/marital group – is also included. 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 1% level, ** = significant at 5% level, * = significant at 10% level

Table 7: Does Marital Status Affect Pregnancy Wantedness?							
		(1) No controls	(2) Plus Mother's Age	(3) Plus Mother's Education	(4) Plus Mother's Race	(5) Plus Previous Births	(6) Plus Insurance Status
Coefficient on "Married?"	15-44 Year Olds	0.380***	0.274***	0.253***	0.251***	0.253***	0.251***
		(0.008)	(0.010)	(0.010)	(0.011)	(0.011)	(0.012)
	N	45,059	45,059	44,264	44,062	43,789	31,290
	20 to 24 Year Olds	0.247***	---	0.256***	0.256***	0.259***	0.284***
		(0.019)		(0.020)	(0.020)	(0.020)	(0.022)
	N	9,594		9,423	9,370	9,314	6,777
	25 to 29 Year Olds	0.323***	---	0.301***	0.300***	0.301***	0.267***
		(0.016)		(0.018)	(0.018)	(0.018)	(0.022)
	N	12,351		12,171	12,113	12,036	8,490
	30 to 34 Year Olds	0.263***	---	0.218***	0.211***	0.206***	0.206***
		(0.022)		(0.023)	(0.023)	(0.023)	(0.027)
	N	11,861		11,659	11,614	11,549	8,179

Notes: Sample drawn from the PRAMS. Standard errors in parentheses. Omitted categories for added controls: (2) over 40 years old, (3) 0-8 years of education, (4) non-white; (6) uninsured. State fixed effects included in all specifications.

*** = significant at 1% level, ** = significant at 5% level, * = significant at 10% level. Weighted.

Table 8: Does Health Insurance Affect Access to Primary Care Physicians?					
		(1) No controls	(2) Plus Income	(3) Plus Employment Status	(4) Plus Age
Coefficient on "Has Health Insurance?"	Entire Sample (N=3,655)	0.251*** (0.030)	0.224*** (0.030)	0.224*** (0.030)	0.221*** (0.031)
	Married (N=1,970)	0.193*** (0.036)	0.160*** (0.035)	0.160*** (0.034)	0.161*** (0.034)
	Single (N=1,685)	0.267*** (0.035)	0.253*** (0.032)	0.253*** (0.032)	0.248*** (0.034)
		(5) Plus Health Status	(6) Plus Disability	(7) Plus Education	(8) Plus Marital Status
Coefficient on "Has Health Insurance?"	Entire Sample (N=3,655)	0.223*** (0.030)	0.221*** (0.030)	0.223*** (0.030)	0.223*** (0.030)
	Married (N=1,970)	0.162*** (0.035)	0.159*** (0.035)	0.159*** (0.035)	---
	Single (N=1,685)	0.250*** (0.032)	0.249*** (0.032)	0.252*** (0.034)	---

Notes: Sample drawn from 2004 BRFSS and is restricted to New England states. Standard errors in parentheses. Regressions (1) – (8) exclude women of childbearing age who are using permanent contraception methods like tied tubes and hysterectomy or whose partner is using permanent contraception methods like vasectomy. Standard errors in parentheses (clustered at state level). Omitted categories for added controls: (2) less than $10,000, (3) unemployed, (5) in bad health; (6) not having a disability, (7) never attended school, (8) single. State fixed effects included in all specifications. For the entire sample, 88.5 percent have access to a primary care physician; among married women it is 91.4 percent and among single women it is 85.0 percent.

*** = significant at 1% level, ** = significant at 5% level, * = significant at 10% level

Table 9: Does Access to Primary Care Physicians Affect Contraception Use?

		Any Contraception Use?				
		(1) No controls	(2) Plus Age	(3) Plus Education	(4) Plus Marital Status	(5) Plus Number of Children
Coefficient on "Access to Primary Care Physician?"	Entire Sample (N=3,655)	-0.025 (0.019)	0.007 (0.018)	0.001 (0.015)	0.009 (0.014)	0.006 (0.015)
	Married (N=1,970)	0.006 (0.014)	0.014 (0.014)	0.005 (0.018)	---	-0.003 (0.022)
	Single (N=1,685)	-0.017 (0.029)	0.019 (0.028)	0.018 (0.026)	---	0.019 (0.026)
		Effective Contraception Use?				
		(1) No controls	(2) Plus Age	(3) Plus Education	(4) Plus Marital Status	(5) Plus Number of Children
Coefficient on "Access to Primary Care Physician?"	Entire Sample (N=3,655)	0.093*** (0.014)	0.132*** (0.008)	0.127*** (0.006)	0.136*** (0.006)	0.136*** (0.007)
	Married (N=1,970)	0.111** (0.028)	0.125*** (0.029)	0.120** (0.030)	---	0.117** (0.031)
	Single (N=1,685)	0.118** (0.037)	0.158*** (0.030)	0.157*** (0.027)	---	0.157*** (0.028)

Notes: Sample drawn from 2004 BRFSS and is restricted to New England states. Standard errors in parentheses. Specifications (1)-(5) exclude women of childbearing age who are using permanent contraception methods like tied tubes and hysterectomy or whose partner is using permanent contraception methods like vasectomy. Standard errors in parentheses (clustered at the state level). Columns (1) through (5) show regressions results with incrementally added controls. Omitted categories for added controls: (3) never attended school (4) single. State fixed effects included in all specifications. For the entire sample, 74.2 percent use contraception; among married women it is 69.4 percent and among single women it is 79.9 percent. For the entire sample, 47.6 percent use effective contraception; among married women it is 41.8 percent and among single women it is 54.4 percent.

*** = significant at 1% level, ** = significant at 5% level, * = significant at 10% level

Appendix Table 1: Summary Statistics

Reported	0.079	0.015	0.128	0.061	0.206	0.056	0.081	0.079	0.080	0.078
Pregnancy	(0.270)	(0.120)	(0.334)	(0.240)	(0.405)	(0.230)	(0.272)	(0.269)	(0.272)	(0.268)
Baby Based	0.075	0.023	0.120	0.058	0.194	0.052	0.079	0.075	0.076	0.072
On HH Roster	(0.264)	(0.149)	(0.325)	(0.233)	(0.395)	(0.221)	(0.270)	(0.264)	(0.265)	(0.259)
Health Insurance	0.918	0.942	0.893	0.856	0.936	0.934	0.906	0.909	0.949	0.907
Coverage	(0.075)	(0.046)	(0.089)	(0.084)	(0.073)	(0.060)	(0.096)	(0.067)	(0.051)	(0.076)
Massachusetts?	0.444	0.440	0.449	0.469	0.424	0.440	1.000	0.000	1.000	0.000
	(0.497)	(0.496)	(0.497)	(0.499)	(0.494)	(0.496)	(0.000)	(0.000)	(0.000)	(0.000)
Post?	0.546	0.570	0.560	0.594	0.520	0.525	0.000	0.000	1.000	1.000
(2008-2011)	(0.498)	(0.495)	(0.496)	(0.491)	(0.500)	(0.499)	(0.000)	(0.000)	(0.000)	(0.000)
Age-Specific	0.083	0.183	0.076	0.093	0.054	0.054	0.072	0.070	0.094	0.092
Unemployment Rate	(0.056)	(0.048)	(0.036)	(0.029)	(0.031)	(0.026)	(0.046)	(0.052)	(0.057)	(0.060)
HS Dropout	0.173	0.716	0.084	0.099	0.066	0.062	0.152	0.157	0.184	0.190
	(0.378)	(0.451)	(0.277)	(0.299)	(0.247)	(0.242)	(0.359)	(0.364)	(0.387)	(0.393)
HS Graduate	0.479	0.284	0.547	0.602	0.482	0.483	0.475	0.537	0.419	0.480
	(0.500)	(0.451)	(0.498)	(0.489)	(0.500)	(0.500)	(0.499)	(0.499)	(0.493)	(0.500)
Did Not Move	0.964	0.970	0.945	0.941	0.951	0.981	0.971	0.964	0.963	0.963
Between States	(0.185)	(0.170)	(0.228)	(0.236)	(0.217)	(0.137)	(0.169)	(0.187)	(0.190)	(0.189)
Military Service	0.008	0.001	0.008	0.007	0.009	0.011	0.006	0.012	0.005	0.009
	(0.090)	(0.029)	(0.089)	(0.081)	(0.097)	(0.104)	(0.078)	(0.107)	(0.074)	(0.092)
Non-Citizen	0.076	0.036	0.094	0.063	0.130	0.074	0.088	0.057	0.098	0.067
	(0.265)	(0.187)	(0.291)	(0.243)	(0.336)	(0.261)	(0.284)	(0.231)	(0.297)	(0.249)
African-American	0.050	0.054	0.057	0.075	0.036	0.042	0.051	0.040	0.059	0.051
	(0.218)	(0.226)	(0.232)	(0.264)	(0.186)	(0.201)	(0.220)	(0.195)	(0.236)	(0.221)
Hispanic	0.101	0.098	0.119	0.113	0.126	0.086	0.103	0.076	0.135	0.094
	(0.302)	(0.298)	(0.324)	(0.317)	(0.332)	(0.280)	(0.304)	(0.265)	(0.341)	(0.293)
Other Non-white	0.078	0.083	0.096	0.108	0.080	0.060	0.068	0.060	0.093	0.088
	(0.268)	(0.276)	(0.294)	(0.311)	(0.271)	(0.237)	(0.252)	(0.238)	(0.290)	(0.283)
Married?	0.529	0.007	0.457	0.000	1.000	0.786	0.550	0.567	0.497	0.510
	(0.499)	(0.085)	(0.498)	(0.000)	(0.000)	(0.410)	(0.498)	(0.495)	(0.500)	(0.500)
Ages 15-19	0.155	1.000	0.000	0.000	0.000	0.000	0.145	0.149	0.161	0.163
	(0.362)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.352)	(0.356)	(0.367)	(0.370)
Ages 20-24	0.121	0.000	0.294	0.478	0.075	0.000	0.115	0.107	0.133	0.128
	(0.326)	(0.000)	(0.455)	(0.500)	(0.263)	(0.000)	(0.319)	(0.310)	(0.339)	(0.334)
Ages 25-29	0.128	0.000	0.309	0.307	0.311	0.000	0.122	0.123	0.134	0.130
	(0.334)	(0.000)	(0.462)	(0.461)	(0.463)	(0.000)	(0.328)	(0.328)	(0.340)	(0.336)
Ages 30-34	0.164	0.000	0.398	0.215	0.614	0.000	0.166	0.169	0.162	0.161
	(0.371)	(0.000)	(0.489)	(0.411)	(0.487)	(0.000)	(0.372)	(0.375)	(0.368)	(0.367)
Age 35-39	0.207	0.000	0.000	0.000	0.000	0.480	0.221	0.218	0.194	0.199
	(0.405)	(0.000)	(0.000)	(0.000)	(0.000)	(0.500)	(0.415)	(0.413)	(0.395)	(0.399)
Age 40-44	0.224	0.000	0.000	0.000	0.000	0.520	0.230	0.234	0.217	0.219
	(0.417)	(0.000)	(0.000)	(0.000)	(0.000)	(0.500)	(0.421)	(0.423)	(0.412)	(0.413)
Under 150%	0.158	0.152	0.204	0.279	0.117	0.115	0.134	0.151	0.157	0.179
of the FPL	(0.364)	(0.359)	(0.403)	(0.448)	(0.321)	(0.319)	(0.341)	(0.358)	(0.364)	(0.383)
Between 150 and	0.135	0.123	0.150	0.146	0.154	0.126	0.119	0.149	0.120	0.147
250% of FPL	(0.342)	(0.329)	(0.357)	(0.353)	(0.361)	(0.332)	(0.324)	(0.356)	(0.325)	(0.354)
Between 250 and	0.075	0.072	0.076	0.070	0.083	0.076	0.072	0.085	0.068	0.076
300% of the FPL	(0.264)	(0.259)	(0.265)	(0.255)	(0.276)	(0.265)	(0.259)	(0.278)	(0.251)	(0.264)
Over 300% of the	0.632	0.653	0.570	0.505	0.646	0.684	0.674	0.616	0.655	0.599
FPL	(0.482)	(0.476)	(0.495)	(0.500)	(0.478)	(0.465)	(0.469)	(0.486)	(0.475)	(0.490)
Sample size	507,000	78,763	209,477	113,701	95,776	218,760	95,051	134,910	129,823	147,216
Sample	All	Age 15-19	Age 20-34	Age 20-34, Unmarr.	Age 20-34, Marr.	Age 35-44	In Mass., Pre	In New Eng., Pre	In Mass., Post	In New Eng., Post

Notes: Sample drawn from the 2003-2011 ACS (excluding 2007). Standard deviations in parentheses.

Figure 1: Expected Fertility Effects by Age, Marital Status and Gains In Insurance Coverage

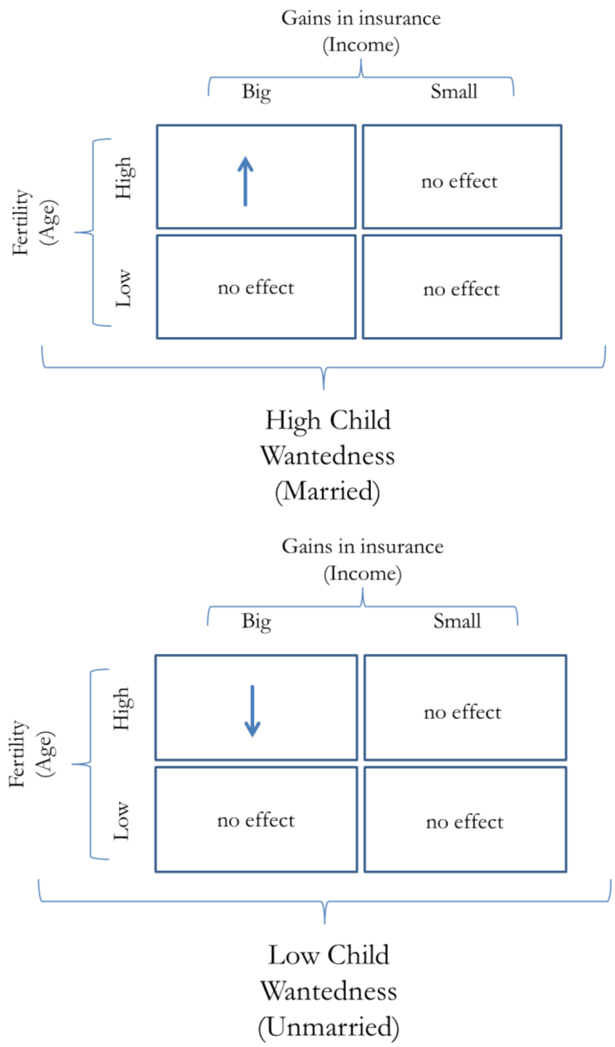


Figure 2: Insurance Coverage Rates By Age Group

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

Figure 3: Insurance Coverage Rates By Marital Status

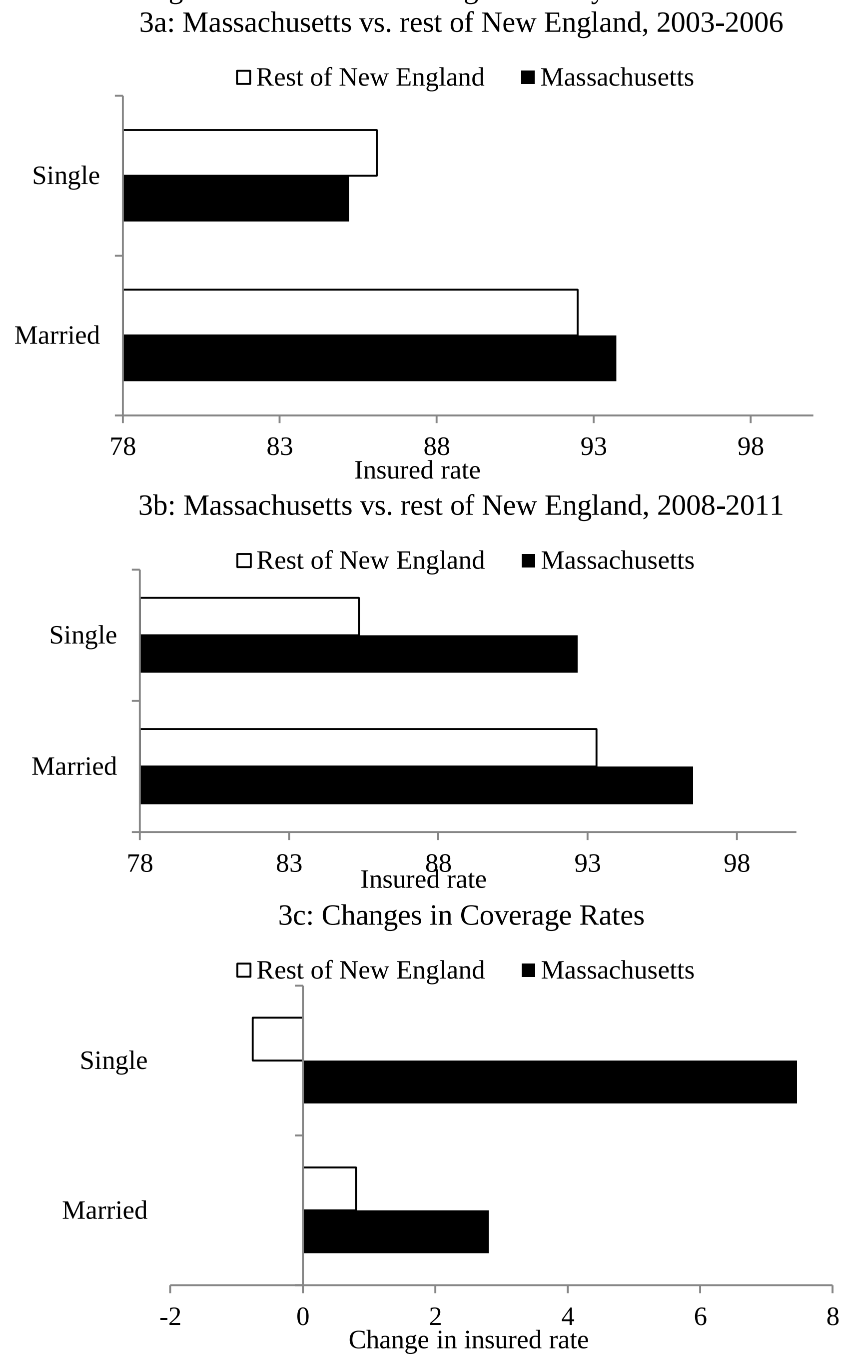


Figure 4: Insurance Coverage Rates By Income Group

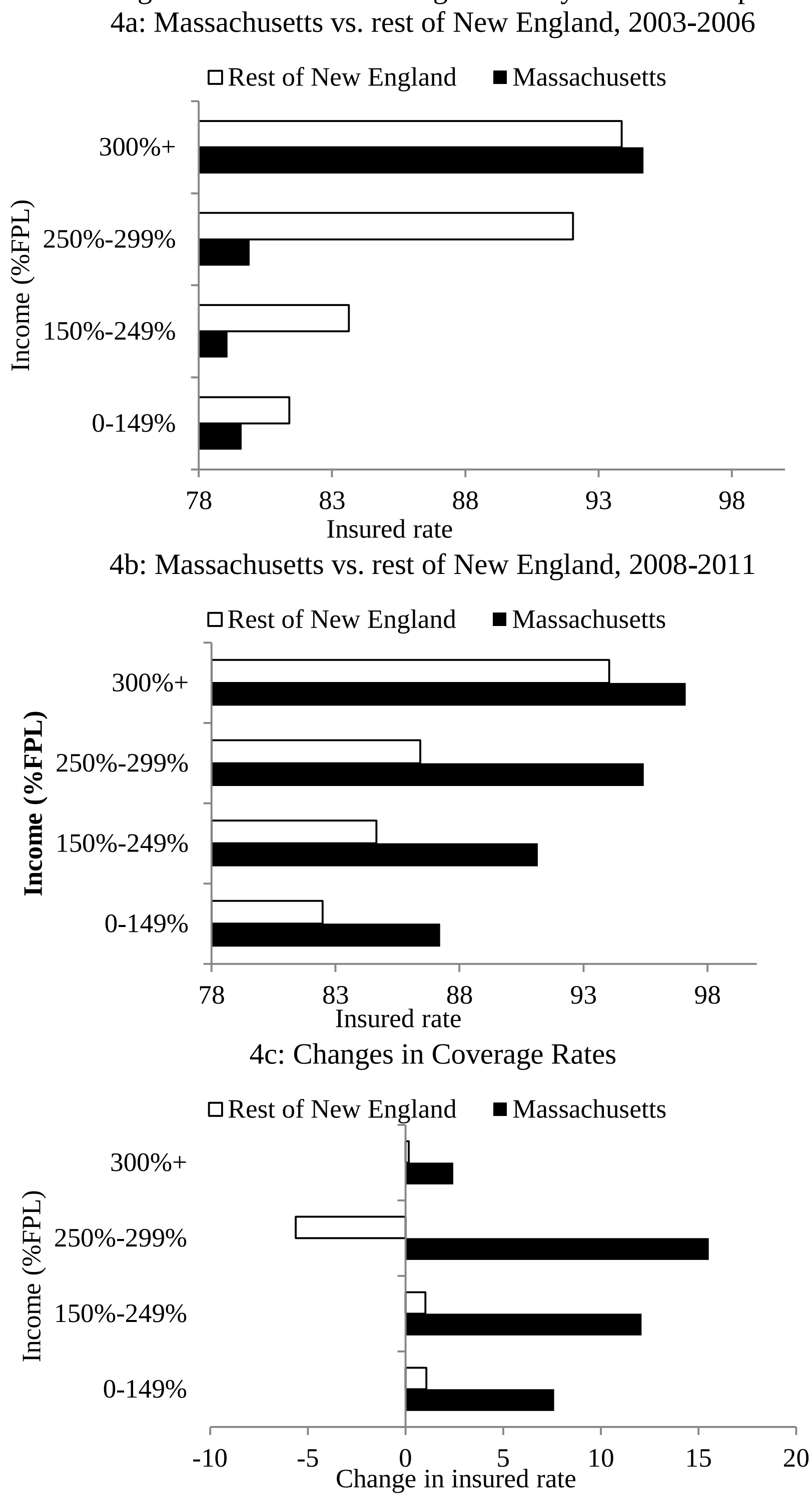
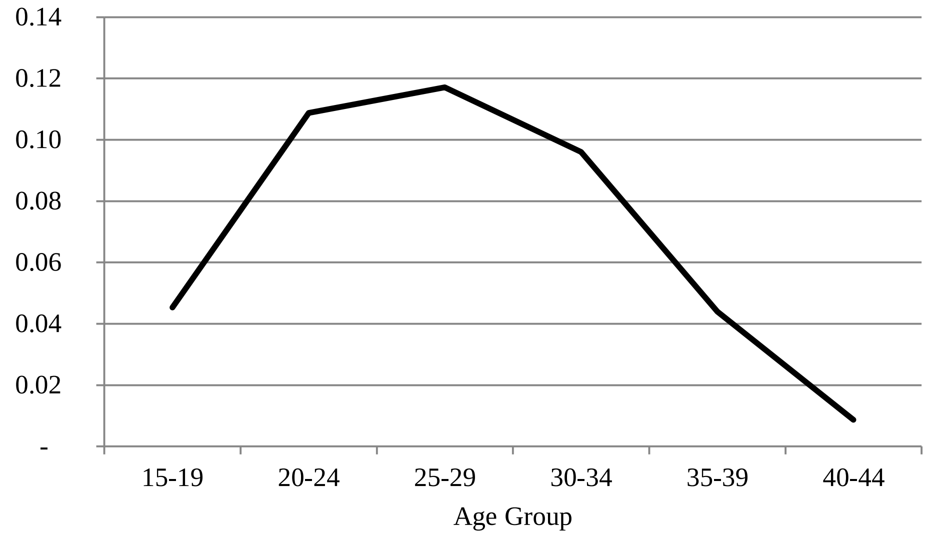


Figure 5: Fertility Rates

5a: Fertility Rates by Age, 2003



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

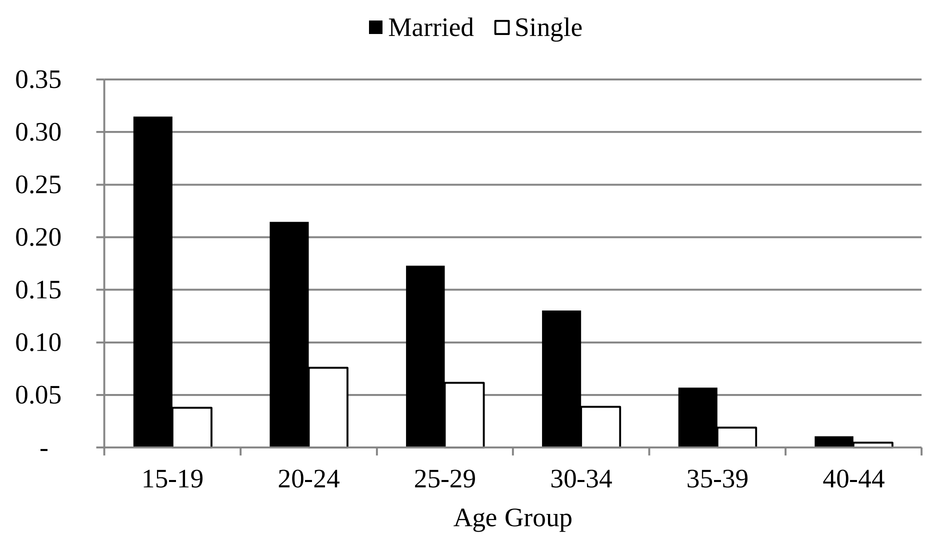
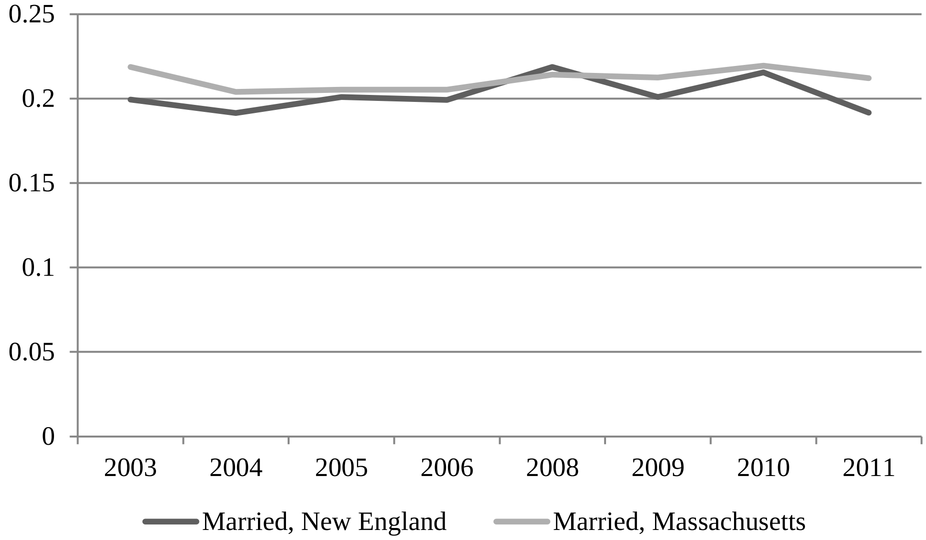


Figure 6: Differential Pre-Existing, Pre-Program Trends in Fertility Rates?

6a: Married Women in ACS, Aged 20-34



6b: Single Women in ACS, Aged 20-34

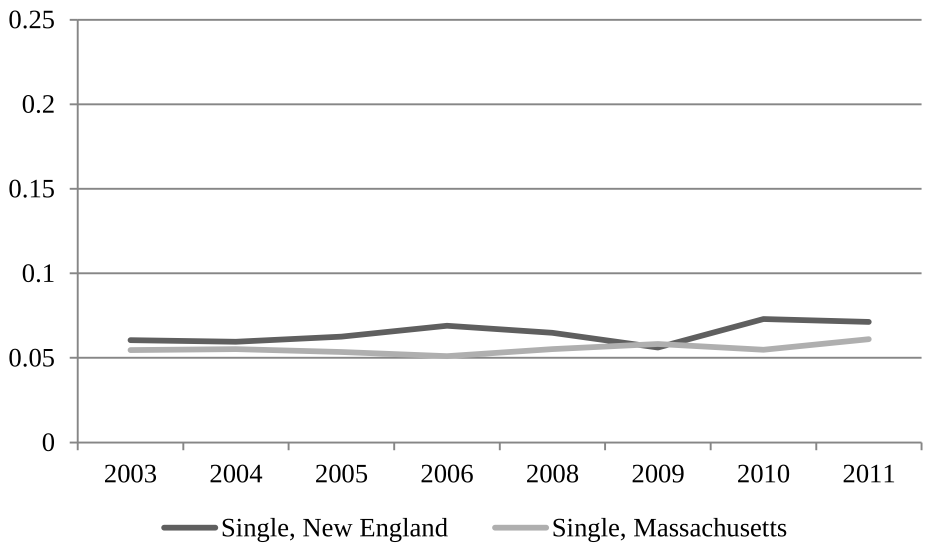
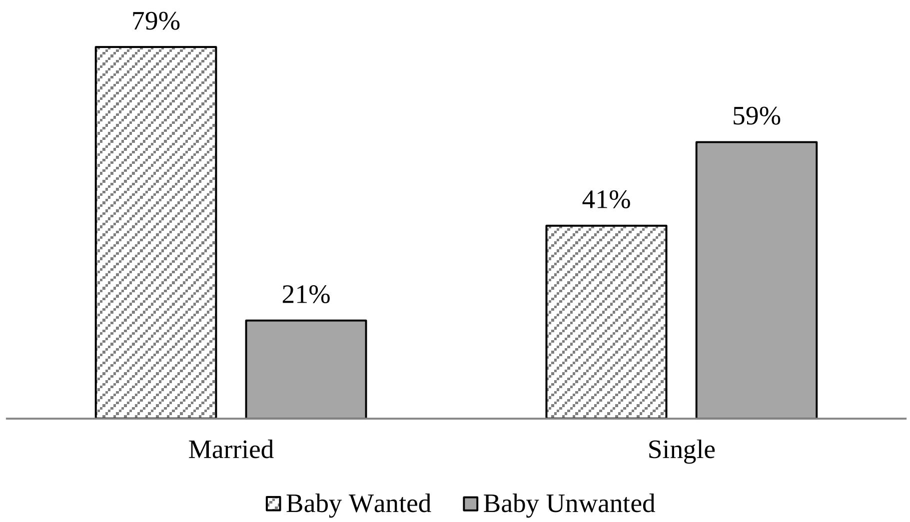
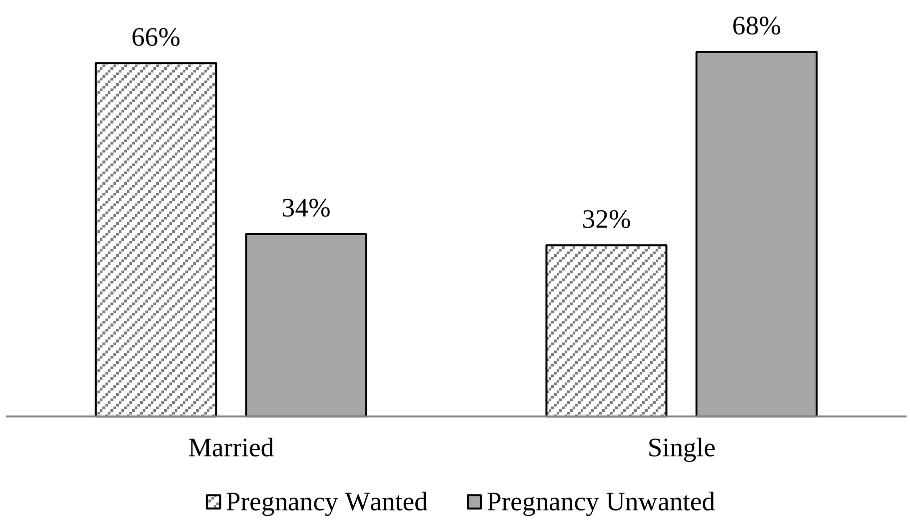


Figure 7: Pregnancy Wantedness by Marital Status

7a. Live Births Only



7b. Live Births and Abortions



Notes: Authors' tabulations using data from the Guttmacher Institute and the PRAMS survey using weighted data for Massachusetts (2007-2011), Vermont (2003-2011), Maine (2003-2011), and Rhode Island (2003-2011).