Do Insurers Risk-Select Against Each Other? Evidence from Medicaid and Implications for Health Reform*

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Abstract

Increasingly in U.S. public insurance programs, the state finances and regulates competing, capitated private health plans but does not itself directly insure beneficiaries through a public fee-for-service (FFS) plan. We study how high- and low-cost groups fare in these settings using county transitions from FFS Medicaid to capitated Medicaid managed care (MMC) for pregnant women and infants. We first document the large health disparities and corresponding cost differences between blacks and Hispanics (who make up the large majority of Medicaid enrollees in our data), with black births costing nearly double that of Hispanics. We find that black-Hispanic infant health disparities widen under MMC—e.g., the infant mortality rate increases by 12 percent percent for blacks while *decreasing* 22 percent for Hispanics—and black mothers' quality of care worsens relative to that of Hispanics. Remarkably, black birth rates fall (and abortions rise) significantly after MMC—consistent with mothers reacting to poor care by reducing fertility or plans discouraging births from high-cost groups. Our empirical findings are consistent with a simple model of risk-selection, where capitation incentivizes insurers to retain low-cost clients and thus improve their care relative to high-cost clients, who they prefer would switch to a competitor. Implications for the ACA exchanges are discussed.

Keywords: risk-selection; Medicaid Managed Care; insurance exchanges; birth outcomes

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1 Introduction

Increasingly in U.S. public insurance programs, the state finances and regulates competing, capitated private insurance plans but does not itself directly insure beneficiaries through a public fee-for-service (FFS) plan. Whereas Medicare debuted in 1965 as a traditional publicly administered FFS program, the 2010 Affordable Care Act (ACA) will expand insurance coverage almost entirely through this new private model. The ACA insurance exchanges offer private, capitated insurance plans with substantial government subsidies and regulation, but no public FFS option. The large majority of the ACA Medicaid expansion will occur under this private model as well, as most states have switched from FFS Medicaid to Medicaid Managed Care (MMC), where Medicaid enrollees choose from private, capitated plans with no FFS option.[1] There is substantial support for similarly changing Medicare—while it currently offers a traditional public FFS plan that competes alongside private, capitated Medicare Advantage (MA) plans, in 2011 the House of Representatives passed a bill eliminating the FFS option.[2]

By 2019, via the exchanges and MMC, nearly sixty million Americans will be covered through this model of strictly private provision of public insurance, yet relatively little evidence exists on how it delivers health care and impacts health outcomes.[3] In particular, a central concern for any insurance program is the possibility that insurers or providers might avoid high-risk individuals (Newhouse, 1996). Yet the existing evidence on risk-selection does not directly apply to the setting we describe above. Much of the research on risk-selection has focused on cream-skimming by private MA plans. These papers generally find that MA

[1]See `http://www.medicaid.gov/Medicaid-CHIP-Program-Information/By-Topics/Data-and-Systems/Downloads/2011-Medicaid-MC-Enrollment-Report.pdf`. In 2011, MMC accounted for 74 percent of all Medicaid enrollees.

[2]The proposal was known as Medicare Premium Support and failed in the Senate.

[3]See `http://www.cbo.gov/sites/default/files/cbofiles/attachments/43472-07-24-2012-CoverageEstimates.pdf`, Table 3. CBO estimates that by 2019, 25 million individuals will be on exchanges and 43 million in Medicaid. As MMC currently accounts for 74 percent of all Medicaid enrollees (see footnote 1), we estimate that 0.74*43 = 32 million will be on MMC. This figure is likely an underestimate as the MMC share of Medicaid enrollees has been steadily growing and will likely exceed 74 percent by 2019. All CBO enrollment projections reflect the June 2012 Supreme Court decision limiting the Medicaid expansion.

plans appear able to direct high-cost beneficiaries to the public FFS Medicare pool, resulting in overpayments to MA plans.[4] But this channel of risk-selection is not available to private plans in settings such as the ACA exchanges or MMC that do not offer a public FFS option. Similarly, many papers have found that without guaranteed-issue or community-rating mandates, private plans risk-select by simply denying coverage to high-cost enrollees or charging them higher premiums.[5] But ACA exchange regulations essentially eliminate this possibility, and MMC plans must accept any eligible enrollee at the same premium ($0).[6]

However, in this paper we posit that risk-selection can still occur in these new "exchange" settings when plans compete against each other and face incentives to retain low-cost individuals while encouraging high-cost enrollees to switch to other plans. We begin by empirically examining the experience of high- and low-cost pregnant women and infants in the MMC setting. For this analysis, we require, first, a setting in which health care provision switches to capitation in a quasi-experimental manner, and, second, groups of enrollees who we, as researchers, can identify *ex-ante* as having different baseline health (and thus expected costs). We argue that Texas' county-by-county transition from FFS to MMC is a suitable experiment and focus on changes in outcomes and care between U.S.-born blacks and Hispanics (and use foreign-born members of these groups, who generally did not qualify for Medicaid during our sample period, as placebo groups). In Texas, these two treatment groups make up the majority of Medicaid births, and, as we document, experience vastly different health outcomes. For example, children of U.S.-born black mothers have seventy percent greater mortality and low-birth-weight rates than children of U.S.-born Hispanic mothers.[7] These disparities translate into enormous differences in costs and profits—in Texas hospital discharge data, black infants have charges that are over eighty percent greater than those of

[4]See, e.g., Langwell and Hadley (1989), Physician Payment Review Commission (1997), Mello *et al.* (2003) and Batata (2004).

[5]See Baicker and Dow (2009), for example, and citations therein.

[6]ACA insurers are subject to guaranteed issue but may modestly adjust premiums for age and smoking status.

[7]This black-Hispanic gap in health has also been widely documented in other settings. We review this evidence in Section 7.

Hispanics, yet MMC plans receive the same capitation payment for the two groups.

Having identified our *ex-ante* high- and low-cost groups, we use administrative birth records data from Texas over 1993-2001 to explore how their outcomes change after the switch from FFS to MMC. We find that MMC leads to a striking increase in inequality in health at birth for blacks relative to Hispanics. Mortality rates for infants born to U.S.-born black mothers significantly increase (by 12 percent) while those for Hispanics significantly *decrease* (by 22 percent), causing the black-Hispanic mortality gap to grow by 61 percent. The black-Hispanic low-birth-weight and pre-term-birth gaps also increase significantly. Infants born to foreign-born black and Hispanic mothers show no such changes in their outcomes after their counties switch to MMC, suggesting that the effects we see for U.S.-born mothers are driven by changes in policy, and not coincident changes in unobserved factors related to health. Our results are robust to different specifications, such as those that include measures of local economic conditions.

We also find that after MMC, black mothers, relative to Hispanics, are less likely to begin prenatal care in their first month of pregnancy and are more likely to receive prenatal care at a public clinic (as opposed to a hospital or private physician's office). These results suggest that black women experience a significant decline in access to care and providers during their pregnancies, while Hispanic women do not.

Finally, we find a change in birth rates among high-cost women after MMC. We document that the switch to MMC leads to a significant decrease in births to U.S.-born black women, concentrated among those who are unmarried. Increased abortions appear to account for a significant share of this effect, though data limitations prevent us from making a more definitive assessment of the role of abortions versus conceptions. In summary, under MMC, infants whose costs are likely to exceed the capitation payment ("high-cost") die more frequently, experience worse health outcomes, receive diminished care, or are not born at all.

Why might MMC increase disparities in care and outcomes between high- and low-cost individuals? We offer a simple model of insurer incentives to explain this new form of risk-selection. We consider two types of patients: high- and low-cost (i.e., those whose expected costs are above and below the capitation payment, respectively). We assume that all patients always have a choice to enroll in one of at least two private plans. This condition is met in our data, generally met by state MMC programs and widely expected (indeed, promised) to be met in the ACA exchanges.[8] As in Hart *et al.* (1997), we also assume that the state cannot write complete contracts with insurers and thus some scope remains for them to adjust quality of care across enrollees.[9] In fact, in our setting, Texas explicitly encourages MMC plans to use their discretion to tailor many benefits individually.

Each period, profit-maximizing plans choose the level of care to provide to the high- and low-cost patients. The probability that a patient returns to the same plan in the following period (e.g., in our context of Medicaid births, that she chooses the same plan for her child's subsequent care or for her next pregnancy) increases with the level of care. Consequently, plans have an incentive to retain low-cost, profitable patients and thus provide them with greater levels of care relative to high-cost patients. By contrast, plans balance two competing incentives in treating a high-cost, unprofitable patient—although reducing the level of care may worsen her outcomes and increase costs in the current period, it will also encourage her to switch to a competitor in the next period. Unlike many models of adverse selection, plans in our framework need not be able to predict the costs of enrollees *ex-ante* or devise a menu of services that encourage the healthy to self-select (though they may also engage in such tactics as well). Instead, they can learn about patient costs and profitability over time and adjust quality of care accordingly based on whether they wish to retain the patient.

Our model has three main predictions. First, relative to FFS, MMC will lead to improved care for healthy, low-cost pregnant women and infants and worse care for unhealthy, high-cost

[8] See `http://kaiserfamilyfoundation.files.wordpress.com/2013/01/8046-02.pdf`, p. 2, on the prevalence of multiple plans in state MMC programs. Unfortunately, the authors write that "generally" enrollees have a choice of at least two plans, but do not give an exact share. Regarding the ACA exchanges, the Obama Administration has claimed that, based on current data, 90 percent of exchange enrollees will have the choice of at least five insurers: `http://www.washingtonpost.com/blogs/wonkblog/wp/2013/05/30/are-obamacares-exchanges-competitive-heres-what-the-experts-say/?wprss=rss_ezra-klein`.

[9] Given that insurers must approve providers' charges, this adjustment in care can be directly passed down to providers.

pregnant women and infants. Second, as a result of the changes in care, health disparities between low-cost and high-cost infants will grow under MMC. Third, birth rates from high-cost groups will fall, either because, as suggested in past work, mothers react to diminished care and outcomes by reducing fertility or because plans specifically discourage births from unprofitable groups.[10]

Our empirical results support the predictions from this model. Importantly, the fact that there are *improvements* for Hispanics argues against an alternative model in which the quality of care generally deteriorates under MMC relative to FFS and those with already poor health fair worse. Such a model would predict either small negative or no effects on Hispanic infant health, not the substantial improvements we find—e.g., the mortality decreases for Hispanic children are actually larger (both in magnitude and when compared to sub-sample means) than the mortality increases for blacks.

We believe our paper makes several contributions. First, it is the only paper, to our knowledge, to examine how high- and low-cost groups fare in a public "exchange" setting—where the government regulates and finances competing private plans but does not itself directly insure individuals. Given the high level of inequality in birth outcomes between blacks and Hispanics before MMC, our finding that the gap in health increases substantially when these groups are under a capitation system is notable and policy-relevant. We argue that, when viewed dynamically, the possibility a client can switch plans means competing plans under capitation are not in fact true residual claimants on their patients' future costs and thus under-provide care to high-cost patients. In settings like Medicare with both private capitated plans and a public FFS option, the government bears the cost of risk-selection in the form of overpayments to private plans; in exchange settings, by contrast, this cost is borne by enrollees with costs above the capitation payment.

Second, while the comparison is not often made, the ACA exchanges will mirror MMC in many ways and thus our results have implications for health care reform. In addition to

[10] On the connection between fertility and health care quality, see, e.g., Albanesi and Olivetti (2010), which we discuss later in the paper.

the similarities already noted, since the ACA exchange subsidies are generally unavailable to those with employer insurance, there will likely be substantial "churn" in the exchange population as people come and go based on their outside options (as in Medicaid). As such, the data collection necessary for effective risk-adjustment in both contexts is difficult, and in fact few states attempt to adjust MMC capitation payments based on *ex-ante* health conditions.[11] We discuss additional challenges common to both settings in Section 8.

Third, our paper further enriches the current understanding of incentives under MMC, which by 2019 will serve as the primary insurer for roughly 32 million individuals (ten percent of the U.S. population) and is thus of interest in its own right.[12] Two important papers have made use of the county-by-county switch from FFS to MMC in California, the same empirical strategy that we adopt for Texas. Duggan (2004) finds MMC increased costs in California, which he attributes to competing MMC plans' limited ability to negotiate favorable rates with providers relative to a consolidated FFS system.[13] Aizer *et al.* (2007) find that pre-natal care and birth outcomes deteriorate under MMC in California.[14] California's MMC model does not provide an ideal setting for an evaluation of the treatment of high- vs. low-cost patients, however, as the program "carves out" expensive patients, passing them back to the state. As health costs are heavily skewed, understanding the experience of high-cost patients under MMC is important for predicting future health care costs. Additionally, comparing our results with past work can help illustrate the tradeoffs involved in MMC program design.

Fourth, this paper finds results along a new margin—*where* pre-natal care is received—that suggests cost-shifting on the part of MMC plans. That women on MMC would turn to

[11]See Winkelman and Damler (2007). They find that only 13 states have implemented MMC risk-adjustment or are considering it. Also see Weiner *et al.* (2012) for a review of issues related to risk-selection in the ACA insurance exchanges.

[12]See footnote 3 for the 32 million figure.

[13]Duggan and Hayford (2011) find supporting evidence that MMC increased costs relative to Medicaid FFS nationally using state panel data.

[14]There is an earlier literature on the effect of MMC on pre-natal care and birth outcomes on which we do not focus. Findings from these papers are mixed perhaps because many rely on cross-sectional comparisons and pre/post analyses without comparison groups (see Kaestner *et al.*, 2002 for an overview). Such research designs may suffer from omitted variables bias due to individual selection into managed care or concurrent macroeconomic trends.

clinics meant for the uninsured suggests that plans may be directing their enrollees to free care instead of covering these costs themselves. In this sense, any savings under MMC are partially offset (or cost increases understated) by cost-shifting toward other safety-net programs, an example of the potential substitutability across social-welfare programs discussed by Borghans *et al.* (2012).

The remainder of the paper is organized as follows. Section 2 describes the transition to MMC in Texas and, in particular, documents the wide discretion that plans have in denying certain benefits to enrollees, adding plausibility to our thesis that they have the scope to improve care for profitable enrollees at the expense of unprofitable ones. Section 3 describes how we select our *ex-ante* high- and low-cost groups based on characteristics observable to researchers (and plans) for the empirical analysis. Section 4 describes our main data sources and empirical strategy. Section 5 presents the results on birth outcomes and quality of care. Section 6 presents the results on birth rates. Section 7 provides a theoretical explanation for the empirical results on birth outcomes, quality of care, and birth rates for high-cost (black) and low-cost (Hispanic) groups that we see in the data with a simple dynamic-programming model of plans' incentives under MMC. Section 8 concludes by discussing welfare and policy implications and recommending areas for future work.

2 Background on Medicaid and the Transition to MMC in Texas

Over the last several decades, most states have switched to a managed care system in their administration of Medicaid. In fact, according to the Kaiser Family Foundation, while only about 10 percent of Medicaid recipients were enrolled in managed care in the early 1990s, over 70 percent are today. In Texas, the legislature voted in 1995 to begin a staggered, state-wide shift from traditional Medicaid FFS to Medicaid managed care after experimenting with a small managed care pilot program in four counties. The Texas Health and Human Services Commission (HHSC) set the order in which counties would switch and details are given in Appendix Table 1. According to HHSC officials, small urban areas switched first because they tended to have well-established healthcare provider networks and it would be easier to work out transitional problems in smaller urban areas as opposed to the largest ones. Larger urban counties switched next, with rural counties switching only in 2012. Our identification strategy uses the variation in the timing of the MMC switch across counties and relies on the assumption that this roll-out is uncorrelated with other time-varying county characteristics. It is thus reassuring that the roll-out schedule was set by a central office and not negotiated by individual counties, and indeed Appendix Table 2 shows that a county's switch to MMC is uncorrelated with changes in its population, per-capita income or unemployment rate. The percentage of the Texas Medicaid population enrolled in the managed care program (called State of Texas Access Reform, or STAR) increased from 2.9 in 1994 to 70.8 in 2009.

Participation in managed care among Medicaid enrollees is mandatory following MMC implementation in each county. Enrollees always have at least three plans in their county from which to choose, and are randomly assigned to one if they do not choose a plan by a given deadline after their county's switch to MMC.[15] Default-enrollment rates are low, with over 83 percent of pregnant women actively choosing their plan, suggesting a woman's past experience or knowledge of plan reputation can affect her choice.[16]

In Texas, as in almost all states, pregnant women and infants under one year of age are eligible for Medicaid if their family incomes fall under 185 percent of the federal poverty line (FPL). While older children face somewhat stricter Medicaid eligibility limits (133 percent of FPL for ages 1-5 and 100 percent of FPL for ages 6-18), adult women are essentially only eligible when pregnant. During our sample period, pregnant women who were undocumented immigrants were not eligible for Medicaid (even though their future U.S.-born child would be) and relied on charity and emergency care.[17]

[15] In the rare cases when a Medicaid-eligible woman shows up at the hospital to deliver without having already chosen an MMC plan, she is randomly assigned a plan to cover the cost of the delivery and care of the infant.

[16] We do not have default rates during our sample period; the 83 percent figure is the current share of pregnant women that actively choose their MMC plan. We are grateful to Stephanie Goodman at Texas HHSC for this information.

[17] See `http://www.coderedtexas.org/files/Report_Chapter04.pdf` for background. Recently, Texas has begun to cover Medicaid prenatal care and other services for expectant mothers who are undocumented

Insurance providers receive a capitation payment for each enrollee based on historical Medicaid costs in the locality. For every woman who gives birth, plans receive a Delivery Supplementary Payment and a newborn premium, which are unadjusted outside of these geographical averages. As expensive births will cost far more than these fixed payments, they thus represent a large loss to plans. When we asked the HHSC about whether these basic capitation payments also applied to very high-cost births (by contrast, California creates a "carve-out" for high-cost Medicaid births and returns them to FFS-type reimbursement), we were told that plans would simply make up these losses on profits from low-cost births: "This average [capitation payment] does include the higher cost deliveries and yes, it would under-pay for those but then again it overpays for others to make up for it."[18]

A final, important point about Texas MMC is that plans are encouraged to tailor benefits for each beneficiary. As noted in Texas HHSC Medicaid documentation:

> Value-added services are additional health care services that an MCO [managed care organization] voluntarily elects to provide to its clients at no additional cost to the state. MCOs offer value-added services to attract clients to sign up with them, including adult dental services and diapers for newborns. *Additional services may be offered to clients on a case-by-case basis at the discretion of the MCO* [emphasis added].[19]

Plans thus have discretion to deny services to some enrollees while providing them to others. Optional services the documentation specifically mentions are in-home visitations and free transport to provider locations, though the above quote suggests extremely wide latitude for other services as well. Given this level of plan discretion, it is plausible that plans may be able to for low-cost patients by improving services for them at the expense of their high-cost counterparts.

immigrants.

[18]Email correspondence with the chief actuary for HHSC (March 30, 2012).

[19]See `http://www.hhsc.state.tx.us/medicaid/reports/PB8/PDF/Chp-6.pdf`, p. 6-7.

3 Selecting *Ex-Ante* High- and Low-Cost Groups

To examine the experiences of high- and low-cost groups under MMC, we need to identify two subgroups of the Medicaid population with substantially divergent health care costs based on *ex-ante* observable characteristics to us as researchers (and to the plans). We turn to Texas hospital discharge data for 1999-2004 that provide information on patient race, ethnicity, and age. As discussed in more detail in Section 4, the majority of black and Hispanic births are covered by Medicaid in Texas, and so we begin by comparing initial hospital charges across these two racial groups in Table 1. In these analyses, we control for county and year fixed effects as capitation payments are adjusted in this manner (though these controls have only a modest effect).[20] Unlike the birth-certificate data that we introduce in the next section, the discharge data do not record mother's place of birth, so we cannot look at costs incurred only by U.S.-born mothers and must instead compare all blacks to all Hispanics. Black newborns incur charges 81 percent greater than do their Hispanic counterparts, or, in absolute terms, an additional $4,218 (col. 1).[21] This absolute difference in initial hospital charges substantially understates the cost differences of black and Hispanic infants, as the elevated medical costs of at-risk births persist well beyond the initial hospital stay.[22]

The differences in costs associated with the mother are also substantial, with black mothers incurring 21 percent greater costs than Hispanics (col. 2). This cost gap is not driven by differences in mothers' ages (col. 3), the only relevant individual-level covariate we have in the discharge data.[23]

[20]Unfortunately, discharge data with county identifiers are only available from the third-quarter of 1999 onward, and as such we cannot use it to compare outcomes before and after a county switched to MMC, since our last group of counties switch in January 1999. Consequently, in light of the results in Section 6, the cost differences from the discharge data that we report in Table 1 might be viewed as lower bounds, as plans may have already discouraged the most high-cost black births by 1999.

[21]Hospital charge data are imperfect measures of the final cost to the insurer as plans negotiate discounts from providers. However, these discounts should not vary by demographic groups, so the comparisons in Table 1 give a good approximation of proportional cost differences.

[22]See, e.g., Tommiska *et al.* (2003) and McCormick *et al.* (1991).

[23]As we cannot link mothers to infants in these data, we cannot control for age-of-mother in the infant regressions in col. (1). Obviously, all newborns are the same age, so we cannot adjust for their own ages as we do for mothers.

Figures 1(a) and 1(b) show the raw (unadjusted for county and year) difference in black and Hispanic births by percentile of the cost distribution. The black-Hispanic cost differences for newborns are positive at every centile, and the median difference is roughly \$250. We censor at the 95$^{\text{th}}$ percentile (\$8,452) as otherwise the graph is extremely compressed: the difference at the 99$^{\text{th}}$ percentile is \$76,341. The differences in delivery charges associated with the mother are relatively constant for all percentiles, outside of some noise in the extreme right tail.

As noted, we would have ideally compared the costs of U.S.-born blacks and Hispanics. The only study we know of that compares newborn hospital costs by race *and* place of birth is Reichman and Kenney (1998). They find that in New Jersey, births to all black mothers have hospital costs 57 percent greater than do births to *all* mothers of Mexican origin (the most relevant group for Texas Hispanics); when both groups are limited to those born in the U.S., this gap increases to 72 percent, suggesting that our comparing all black and Hispanic births understates the cost differences between the U.S.-born subsets. On the other hand, as we note in footnote 25, other studies have documented that Mexican-American health outcomes are slightly worse for the generation born in the U.S. relative to the one born in Mexico.

On the whole, Table 1 suggests that blacks and Hispanics may serve as good proxy groups for high- and low-cost Medicaid patients, and so we focus on differences between them throughout our empirical analysis. However, Table 1 suggests that differences by mother's age are also significant. Although we cannot link mothers and newborns so are unable to split newborn costs by the age of the mother, we find that black and Hispanic mothers age 35 and older have delivery costs 15 percent greater and 14 percent greater than their younger counterparts, respectively (see cols. 4 and 5). In our birth certificate data that we describe in the next section, among black and Hispanic mothers the rates of pre-term and low-birth-weight are roughly thirty percent higher for those above age 35. As such, births to older mothers represent a small but expensive subset of Medicaid births that plans would have an incentive to avoid. However, given that older mothers are unlikely to have a future birth, plans may be less worried about these unprofitable clients returning in future periods. For this reason, the empirical work generally focuses on differences by race, though we show some results by mother's age as well.

Before moving on to the empirical work, we emphasize two points. First, the large cost differences between blacks and Hispanics that we document are completely in line with past work on differences in birth outcomes. Hispanic infants in the U.S. are remarkably healthy, and in fact researchers have coined the term "Hispanic paradox" to describe the fact that despite socio-economic deprivation comparable to blacks, they in fact have birth outcomes (and, while less relevant for us, adult health outcomes) equal to or better than non-Hispanic whites.[24] We generally take the cost differences between blacks and Hispanics as given, though briefly review potential explanations in the footnote below.[25]

Second, none of our results that follow shows that plans specifically discriminate against African-Americans *per se*. Rather, we, as researchers, need to use a proxy (race) for *ex-ante* differences in expected medical costs to study how high- and low-cost patients fare in a public "exchange" setting like MMC. If plans choose to tailor care so that they attract low-cost patients while encouraging high-cost patients to switch to a competitor, they may use these proxies as well, or, because they have access to actual cost data throughout the pregnancy, may not need to use the proxies that we do. On the other hand, if plans compete for patients by creating positive (negative) word-of-mouth among low- (high-) cost groups, it might well

[24]See, for example, Leslie *et al.* (2003), Haywood L Brown and Howard (2007), Alexander *et al.* (2003), and Dominguez (2008).

[25]Our reading of the literature suggests that the Hispanic paradox is best explained by the superiority of diet and other health habits in Latin American countries relative to the U.S. (tellingly, these advantage appear to dissipate slightly in the second generation with assimilation, see Guendelman and Abrams, 1995 as well as our Table 2). Another explanation is the "healthy migrant effect"—that only the healthier members of a home country choose to migrate—though Rubalcava *et al.* (2008) find only weak evidence that Mexicans who move to the US are healthier than their counterparts who remain. We do not agree with the claim that the Hispanic paradox is a statistical illusion driven by so-called "salmon bias"—that immigrants wish to return to their home country to die and thus are rendered "statically immortal" in U.S. vital statistics data. First, salmon-bias cannot explain differences in *birth* outcomes. Second, Abraido-Lanza *et al.* (1999) shows that the Hispanic paradox exists for Cuban immigrants (who, for political reasons, almost never return to their home country) and Puerto Rican immigrants (deaths in Puerto Rico are recorded in U.S. vital statistics data).

be profitable for them to target race specifically.

4 Data and empirical strategy

4.1 Main data source

Our main source of data is the universe of birth records from the Texas Department of State Health Services (DSHS). These data contain detailed information on the child's exact birth date, birth outcomes, medical procedures, maternal demographics and health, and the mother's county of residence and country of birth. Using recorded information on each child's birth date and gestation length, we calculate an approximate conception date for each observation. We merge the birth records data to data on the timing of MMC implementation by the mother's county of residence.

As Appendix Table 1 documents, counties switched from FFS to MMC between 1993 and 2006. We drop the four pilot counties that switched in 1993 as we could not determine when the pilot period ended. Since much of the focus of our analysis is on racial differences in birth outcomes and the racial composition of births, we do not analyze the January 2006 switch into MMC as many Texas counties were affected by an influx of black refugees following Hurricane Katrina in September of 2005.[26] Therefore, we limit our sample of analysis to conceptions by mothers residing in Texas between January 1993 and December 2001, allowing for roughly three years before the first MMC switch (in December 1995) and three years after the last MMC switch (in January 1999). Finally, we drop the small number of birth records with missing information on gestation, parity, mother's age, mother's race/ethnicity, and mother's marital status. These sample restrictions leave us with 2,814,681 observations in our main analysis sample.

In Table 2, we present summary statistics for the entire sample, as well as several demographic subsets of mothers: U.S.-born blacks and U.S.-born Hispanics and foreign-born blacks

[26]Results are very similar when we do use the longer sample period and treat the 2006 transition as we do the earlier transitions, and in fact earlier versions of the paper included them before we realized Katrina could contaminate our results. It seems prudent to exclude this transition, however, as several of the counties that switch in 2006 are close to the Louisiana border.

and Hispanics and (all) married white non-Hispanics U.S.-born black and Hispanic mothers are slightly younger than average, and considerably younger than married non-Hispanic white mothers. We compare U.S.-born to foreign born because the former are eligible for Medicaid while the latter are not.

Pre-natal care measures are substantially different for minorities and non-Hispanic whites. Only one-fifth of U.S.-born blacks and Hispanics receive pre-natal care in the first month of pregnancy, whereas 30 percent of married whites do. Whereas less than four percent of married whites receive their pre-natal care in public clinics, 13 and 19 percent of blacks and Hispanics do, respectively.

Differences in U.S.-born black rates of low-birth-weight, pre-term delivery and child death are, respectively, 71 percent, 41 percent, and 74 percent greater than the corresponding rates for U.S.-born Hispanics. Outcomes for infants born to foreign-born Hispanics are slightly better than those born to U.S.-born Hispanics, consistent with the hypothesis that part of the "Hispanic paradox" arises from healthier habits acquired in the home country and erodes slightly with assimilation.

Overall, Table 2 further confirms that U.S.-born blacks and Hispanics may serve as good proxy groups for high- and low-cost Medicaid patients.

4.2 Medicaid coverage

One variable not included in Table 2 is Medicaid status, as Texas only began including it in the birth records data in 2005. However, Medicaid status is actually a problematic variable in the context of studying MMC because privatizing Medicaid has the effect of making enrollees or providers incorrectly record the birth as being covered by a private insurer or by "other/unknown" source, a possibility hypothesized by Aizer *et al.* (2007).

Indeed, we find that this effect is empirically important when we examine Medicaid status for conceptions in 2004-2005 versus 2007-2008 in the counties that privatized in 2006. In the earlier period, 64.7 percent of births in these counties were recorded as covered by Medicaid, compared to 49.9 percent in the later period, suggesting the true Medicaid share is roughly

1.3 times (64.7/49.9) the reported share in the post-period. Similarly, in 2005, the official count of Medicaid births from the Texas DHHS is 1.3 times the count in the birth certificate data.[27] As such, in Appendix Table 3 we gross up by 1.3 the share of Medicaid births in the birth certificate data to estimate the Medicaid share of births for selected groups conceived in 2004 (born in 2005). This correction is of course rough, as different groups might differentially misreport Medicaid coverage under MMC.

Appendix Table 3 reflects the high share of Medicaid-covered births in Texas (in the 2000s, it typically ranked in the top six to eight states in this category, reflecting its relatively low rates of private insurance coverage).[28] Medicaid covered 84 and 88 percent of births to U.S.-born black and Hispanic mothers, respectively, and for Hispanics this share falls only modestly for married mothers (69 percent). Moreover, U.S.-born black and Hispanic mothers account for 56 percent of reported Medicaid births.[29]

The Medicaid share for foreign-born black and Hispanic mothers is far smaller, especially for Hispanics, where the foreign-born Medicaid rate is only thirty percent of the U.S.-born. As noted earlier, during our sample period, undocumented immigrants did not quality for Medicaid. Using estimates of the Hispanic share of the undocumented population from the national level (76 percent, which is certainly a large underestimate for Texas), in the Appendix we estimate a generous lower bound on the share of Texas Hispanic foreign-born mothers in 2000 who are undocumented (and thus ineligible for Medicaid) of 56 percent. The ineligibility of undocumented mothers and the large share of the foreign-born who are undocumented make the foreign-born an excellent placebo group. We will also use married white women in the same manner.[30]

[27] See `http://www.hhsc.state.tx.us/medicaid/reports/PB8/PDF/Chp-4.pdf`, p. 4-15. The official count indicates that 54 percent of births are covered by Medicaid, whereas our birth certificate data indicate 41 percent.

[28] See `http://www.statehealthfacts.org/comparemaptable.jsp?yr=58&typ=2&ind=223&cat=4&sub=57`.

[29] From Appendix Table 3: $(0.836 * 26,615 + 0.877 * 69,146)/(0.539 * 273,471) = 0.562$.

[30] We also have educational attainment in the births data and we could conceivably use it in the creation of placebo groups, but we are uncertain of its accuracy. For example, even among women old enough to have completed high school (age 24 and older, say), women without a high school degree have lower Medicaid rates than those with a high school degree but no college. However, when we use college-educated white

4.3 Empirical Design

In order to assess the causal effects of the switch from FFS to MMC, we exploit variation in the timing of the MMC rollout across counties to create an event-study design. To ease the computational burden, we generally collapse data into county/conception-month cells and weight by cell size (equivalent to estimating the corresponding individual-level regression with no individual-level controls).

Our estimating equation thus takes the form:

$$Y_{ymc} = \beta\, MMC_{ymc} + \Lambda' W_{ymc} + \mu_c + \gamma_y + \nu_m + \mu_c * t + \epsilon_{ymc} \tag{1}$$

for births in county c, conceived in year y, month m. Y_{ymc} is a birth outcome of interest, such as mortality, birth weight or gestation length. MMC_{ymc} indicates that the conception occurred after MMC rollout in county c. W_{yc} is a set of county-year specific controls including population, average income, and the unemployment rate. μ_c are county fixed effects, γ_y are conception-year fixed effects, ν_m are conception-month fixed effects, $\mu_c * t$ are county-specific linear time trends (to follow Aizer *et al.* (2007)), and ϵ_{ymc} is the error term, which we cluster by county. The key coefficient is β, which measures the effect of being conceived under MMC on the outcome of interest. To avoid imposing constraints on coefficients, we estimate equations separately for blacks and Hispanics and then test whether $\beta^{Black} - \beta^{Hisp}$ is of the expected sign and statistically significant. For ease of exposition, unless otherwise noted, "black" and "Hispanic" will refer to U.S-born blacks and U.S.-born Hispanics, respectively.

5 Results on birth outcomes

5.1 Main results

Table 3 compares changes in mortality for black and Hispanics after MMC. For this and many other tables in this section, each pair of columns presents first the estimate for blacks

women as a placebo group instead of married white women, the results are similar and are available upon request. Another reason we focus on marital status and not education is that the abortions data that we introduce in Section 6 do not include information on education.

and then the estimate for Hispanics. Toward the bottom of the table, the "Diff/p-val" row shows in the odd-numbered columns the corresponding differences in the MMC coefficients ($\beta^{Black} - \beta^{Hispanic}$) and in the even-numbered columns shows the p-value associated with the test of equality across the two coefficients.[31]

Cols. (1) and (2) show that mortality—measured by whether a death certificate can be matched with the birth certificate—increases by 0.139 percentage points or ${}^{0.139}/{}_{1.198} = 11.6$ percent among black mothers, while falling by 0.154 percentage points or ${}^{0.154}/{}_{0.715} = 21.6$ percent among Hispanic mothers. Both effects are statistically significant. This 0.293 percentage-point (or ${}^{0.293}/{}_{(1.198-0.715)} = 60.7$ percent) increase in the black-Hispanic mortality gap is itself highly significant ($p < .001$). Adding basic county-time controls for population, income and unemployment has no effect on the results (cols. 3 and 4), not surprising given the earlier result that these trends were uncorrelated with MMC rollout.[32] Cols. (5) and (6) show that for both groups, the magnitude of the effect increases (markedly so for blacks) when only unmarried mothers are included. The gap remains highly significant and indicates that, among births to unmarried mothers, the black-Hispanic mortality gap grows by ${}^{0.429}/{}_{(1.26-0.822)} = 97.9$ percent.

We display these results graphically by substituting the MMC indicator in equation (1) with dummy variables for the 36 months before and after county MMC implementation (normalizing the month of implementation to zero) and plotting these coefficients in Figure 2. Consistent with the regression results, blacks show a positive shift in mortality for children conceived under MMC while Hispanics show a similarly marked, but negative, shift. For both, the shift is coincident with MMC's introduction in a county.

Table 4 shows results for other standard birth outcomes. Again, health significantly worsens for black infants (cols. 1, 3, 5): incidence of pre-term birth (defined, as in the medical

[31] We test equality using seemingly-unrelated regression in Stata, equivalent to running a single regression in which every covariate is interacted with a dummy variable for race.

[32] We use county-year data on per-capita income and population from the Regional Economic Information System (REIS), and unemployment data from the Local Area Unemployment Statistics (LAUS) of the Bureau of Labor Statistics. We interpolate to create monthly measures to avoid sharp jumps at the end of calendar years.

literature, as gestation less than 37 weeks), low birth weight (birth weight less than 2,500 grams), and abnormal birth weight (birth weight less than 2,500 g or more than 4,000 g) increases by 7.5, 5.5 and 6.4 percent, respectively. We also add the sex ratio as an outcome (col. 9), given the growing literature documenting its positive correlation with maternal well-being during pregnancy (as male fetuses are more likely to miscarry due to stress).[33] This outcome moves in the expected direction for black mothers, but is not significant.

The even-numbered columns showing the Hispanic results tell a very different story. While results for birthweight are not significant, the pre-term share falls by 6.3 percent and the male share increases by 0.8 percentage points. For all outcome variables, the change in the black-Hispanic gap is of the expected sign and significant at the five percent level.

Figures 3 and 4 show graphically the results from Table 4 that showed the largest black-Hispanic divergences, the pre-term and male share of births. As with mortality, the divergence in the pre-term share for blacks and Hispanics begins just as a county switches to MMC. The increase in the Hispanic male share also takes place at the time of MMC's introduction (the corresponding effect for blacks is noisier, reflecting the insignificant coefficient in Table 4).

5.2 Robustness checks

Results for placebo groups. As noted in the previous section, foreign-born black and Hispanic mothers have Medicaid coverage rates roughly one-third of their U.S.-born counterparts. Appendix Tables 4 and 5 show that deterioration of U.S.-born black outcomes and improvement of U.S.-born Hispanic outcomes do not extend to the foreign-born population. The only two (marginally) significant outcomes for foreign-born blacks are in the opposite direction of our main U.S.-born results. For foreign-born Hispanics—for whom the sample size is just as large as the U.S.-born subset and we thus have the power to reject smaller effects—the results are all small in magnitude, statistically insignificant and often of the opposite sign of the main U.S.-born results. Appendix Table 6 shows no effect of MMC on

[33] See Fukuda *et al.*, 1998 and Catalano *et al.*, 2005

outcomes for married whites.

Changes in selection. Our results are consistent with blacks receiving relatively lower-quality care relative to Hispanics after MMC, but also with negative changes in selection into birth for black infants versus Hispanics. Appendix Table 5 examines whether incidence of maternal risk-factors differentially change for U.S.-born blacks and Hispanics after MMC. Cols. (1) and (2) show that after MMC mothers in both groups are younger on average (a result we return to later); cols. (3) and (4) show they are both less likely to have diabetes or hypertension, though for neither group is the effect significant; cols. (5) and (6) show blacks are less likely and Hispanics more likely to smoke (though neither result is significant on its own). Of the three outcomes, only one (smoking) show statistically significant black-Hispanic divergences, in the direction of blacks being relatively *positively* selected after MMC, suggesting the effect of MMC on the divergence of birth outcomes in Table 3 and 4 is understated.

Indeed, when we re-run regressions in Appendix Tables 7 and 8 for each of the birth outcome variables using individual-level data and controlling for all plausible pre-determined covariates on the birth certificate (see Table notes), the results are essentially unchanged. Texas no longer provides researchers data with mother identifiers, so we cannot compare siblings born before and after MMC, but given how stable the coefficients are with and without controls and the results on selection in Table 5, we are confident that our birth outcome results are not driven by selection.[34]

Plausibility of magnitudes. The relative effects we find for blacks and Hispanics—especially for mortality, pre-term birth and the sex ratio—are large, but not out of step with past research on the effect of health care on birth outcomes. Aizer *et al.* (2007)'s estimate of the effect of MMC on neonatal death (a fifty percent increase) is larger than the mortality increases we find for black mothers (11.6 percent). As we do for black mothers, Aizer *et al.*

[34]It is also reassuring that Aizer *et al.* find that regressions with and without mother fixed effects yield similar results.

(2007) find larger effects (in percentage terms) on mortality than on pre-term or low-birth weight shares.

Recent work has tied maternal stress and well-being to changes in pre-term birth and the sex ratio. Lauderdale (2006) uses California birth certificate data to show that women with "Arabic-sounding" names exhibited a fifty percent increase in pre-term births after September 11, 2001, presumably as a result of the stress related to the sudden prejudice they experienced. Both acute stress (see, e.g., Fukuda *et al.*, 1998, examining the Kobe earthquake in Japan) and longer-term stress (see, e.g., Catalano *et al.*, 2005, examining unemployment rates) have been tied to changes in the sex ratio or male neonate death rate slightly larger than the effects we documented in Table 4.

5.3 Results on birth inputs

The birth certificate data provide information on pre-natal care and procedures used at birth, though as Reichman and Schwartz-Soicher (2007) document, pre-natal care information on birth certificates (relying on mothers' recall) is less accurate than birth outcomes data.

Table 6 shows results for indicators of pre-natal care. U.S.-born blacks are less likely to receive immediate (within the first month of pregnancy) pre-natal care relative to Hispanics after MMC, though there is no significant difference when later thresholds are used (not shown). There is no difference in the total number of pre-natal visits, though blacks are less likely than Hispanics to receive at least eight visits (we choose this cut-off because almost all women receive at least a handful of visits so there is little variation, whereas visits beyond this point become endogenous to gestation and mother's health). Most women are recommended to gain between 25 and 35 pounds during pregnancy. We choose cut-offs lower than these, as for overweight women less weight gain is recommended and we do not know women's pre-pregnancy BMI. After MMC, black women are more likely than Hispanics to gain insufficient weight during pregnancy, which increases the probability of an infant being small for gestation age and infant mortality (Park *et al.*, 2011; Tenovuo, 1988; Giapros *et al.*, 2012).

The most striking change in pre-natal inputs we found was a shift in the share of black mothers seeking pre-natal care in a public clinic, as opposed to a hospital or private doctor's office. Likely due to the way the question is asked (asking at the time of birth if a women had ever received pre-natal care in a clinic), the effect shows up for those *born* under MMC. As Figure 5 shows, the effect is very large. Table 7 shows that the shift to public clinics among blacks is statistically significant both on its own and relative to Hispanics. Black mothers also show a significant shift away from receiving care in a hospital.

Texas HHSC told us that these "public clinics" would include Planned Parenthood along with "clinics that serve the uninsured." That black mothers would turn to these clinics at such higher rates after MMC is consistent with plans being slower to enroll them or providing them more limited care and access. It would also be consistent with plans contracting with fewer providers in black neighborhoods. Alternatively, MMC plans may simply be directing clients to free clinics as a way to cut their costs regardless of the underlying cost or racial background of the clients. While the coefficient in the public clinics regression is not significant for Hispanics, it is positive and economically non-trivial. Of course, this cost-cutting is merely passing costs on to the state, counties and other parties who fund these clinics. If similar practices occur in other states, then the cost increases that Duggan (2004) attributes to MMC might in fact be understated.

In sum, our examination of results on birth outcomes, selection, and care access indicate that widening health disparities between black and Hispanic births are likely driven by care, rather than selection. As documented in Section 2, MMC plans have vast discretion to tailor services on a case-by-case basis and appear to have directed more attention to lower-cost clients at the expense of higher-cost clients. It is worth noting that we cannot observe many inputs (e.g., the in-home visits or free transportation noted in the Medicaid documentation) and these unobservables might also play a large role in the outcome divergences we document. However, the inputs we can observe generally move in the direction consistent with a significant widening of black-Hispanic health disparities.

6 Results on fertility

Here we test whether births rates of high-cost groups fall after MMC, either because women respond to worse care and outcomes by reducing fertility or because plans actively discourage these women from having children so as to limit their exposure to unprofitable births. We emphasize upfront that we have limited ability to distinguish between these two mechanisms.

6.1 Main results

Table 8 presents regression results based on equation (1) with the share of births to U.S.-born black mothers in a conception county-month now the outcome variable. In col. (1), this share falls by 0.12 percentage points, or 1.1 percent from the sample mean, and does not attain statistical significance. As with the mortality results, however, the fertility effect is concentrated among unmarried black mothers—col. (2) shows that the share of births to this group falls by 0.247 percentage points, or 3.6 percent, and is highly significant.

The dependent variable so far in Table 8 is the *ratio* of black (or unmarried black) births to all births, and thus could conceivably be driven by an increase in non-black births. Col. (3) and (4) regress the log of births to U.S.-born black women and black unmarried women, respectively, and show that the ratio is indeed being driven by a decrease in the numerator of the ratios, with little effect on all other births (col. 5).[35]

Figure 6 (a) shows results graphically, again plotting coefficients from dummies corresponding to 36 months before and after a county switched to MMC. A decrease in level and break from trend coincident with MMC's introduction can be seen. Because of the rapid rise in the Hispanic share of births over our sample period (about 0.83 percentage points per year) and the resulting mechanical decrease in other groups' share, our results could be sensitive to how county time trends are specified. In fact, dropping country trends (subfigure b) or using quadratic trends (subfigure c) produces a more pronounced decrease in the black

[35]We restrict the sample in cols. (4) and (5) to those counties with at least one black unmarried birth in every month, to avoid having to take the log of zeros. These counties account for 87 percent of black unmarried births and 67 percent of all births.

share of births (and the corresponding regression coefficients are of greater magnitude than that in col. 1 using linear trends, not shown), suggesting our choice of linear county time trends is conservative.

As a robustness check, we dropped each county individually to ensure that no single county was driving our results—in all cases, the coefficient in col. (2) remained negative and statistically significant. We also examined whether the "echo" of this birth composition result could be seen in 2005-2011 American Community Survey data as the children born during our sample grew older. We would expect this result to be much noisier, as we proxy county of birth with current county of residence, can only observe quarter instead of month of birth, and can only identify 24 counties as the ACS does not specify smaller counties for confidentiality reasons. Nonetheless, Appendix Figure 1 shows some evidence of a decrease in the black share of cohorts conceived after MMC (though the corresponding regression result is not quite significant at the ten-percent level, results available upon request).

Finally, we note that the decrease in black births is not a result of changing migration patterns among black Texans, either in response to MMC's introduction or merely in a manner coincidentally correlated with it. If our birth rate results were driven by entire families moving we should see enrollment for *school-age* black children decrease when a county switches to MMC. Using administrative data from the National Center for Education Statistics in Appendix Table 9, we find no such effect—the coefficient on MMC introduction is close to zero and switches signs depending on the specification.

6.2 Exploring mechanisms for the fall in black births

Miscarriages. As we have documented in the previous section, the sex ratio has been tied to maternal well-being because male neonates are more likely to miscarry. As such, the decrease in the black sex ratio relative to Hispanics may reflect an increase in the black miscarriage rate and miscarriages might thus explain some of our "missing" black infants. As noted earlier, the Kobe earthquake in Japan led to a 1.5 percentage point decrease in the sex ratio, which was accompanied by a six percent decrease in fertility. Assuming that the decrease in fertility came only through miscarriages (as opposed to fewer conceptions), would suggest our entire effect could be explained by miscarriages, as in Table 4, blacks' sex ratios fell 1.41 percentage-points relative to Hispanics post-MMC. However, such an assumption likely overestimates the share of the post-Kobe fertility effect explained by miscarriages and as far as we know there exists no work that attempts to find the elasticity of the miscarriage rate with respect to the sex ratio at birth.

Abortions. We examine abortion rates from 1998 to 2002, as individual-level abortion data has only been released consistently from 1998 onward. In these regressions, we consider someone as "treated" if they conceive after MMC or within three months of its introduction, since the large majority of abortions in Texas during our sample period occur two to three months after conception.

Given the shorter sample period, this analysis is obviously limited by being identified from a set of counties that all switch at the same time—only those counties that switch in 1999 contribute to the identification—as well as having a relatively short pre-period. Moreover, unlike the birth certificate data, the abortion data do not include place of birth, so we cannot compare foreign- and U.S.-born women.

Nonetheless, the results are striking. In col. (1) of Table 9, the abortion rate among blacks increases by 1.89 percentage points, or 6.7 percent from a sample mean of 27.9 percent.[36] Cols. (2) and (3) show that this effect is driven by unmarried black mothers, consistent with our previous result that the fall in black birth rates is concentrated among unmarried women. Cols. (4) and (5) show, respectively, that the result is unchanged by starting the sample two months later (in case reporting in the first few months of 1998 was less reliable) and is smaller but still significant if county trends are dropped (in case the short pre-period leads to mis-estimated county trends). There is no effect for either Hispanic, Hispanic unmarried mothers or white married mothers (cols. 6, 7 and 8).

[36] Some readers have commented that the mean abortion rates in Table 9 appear too high. In fact, abortion rates in Texas during our sample period were slightly lower than the national average. In 1999, the ratio of abortions to births plus abortions was 18 percent in Texas, whereas the CDC reported that nationally that same ratio was 20.3 percent. See `http://www.cdc.gov/mmwr/preview/mmwrhtml/ss5109a1.htm`.

Re-estimating the main birth-composition result (col. 1 of Table 8) for the same conception-months as in Table 9 suggests that abortion explains 37 percent of "missing" black births. While the short pre-period concerns us, if the effect for unmarried blacks is driven by, say, MMC counties differently improving their abortion reporting over the sample period, we would expect to see some increase as well among similar groups (married blacks, unmarried Hispanics). We thus conclude that abortion appears to be an important factor explaining the fall in black births.

In Texas, Medicaid cannot pay for clients' abortions. However, no restrictions exist on discussing abortion with Medicaid clients and MMC may contract with doctors who are more willing to discuss the option.[37] In addition, it is also possible that women who are on the margin of aborting a pregnancy might be influenced by the initial interactions they have with their providers and insurers. Diminished quality and access to care may make mothers feel that the pregnancy will be a stressful and overwhelming experience, and we showed earlier that care appears diminished for black mothers under MMC.

The result on abortion rates may relate to the increases in pre-natal care provided by public health clinics discussed above, in two potentially reinforcing ways. On the one hand, as noted above, women who receive their pre-natal care at clinics instead of alternative settings may come into contact with providers and other clients who are more likely to suggest abortion as an option. On the other hand, the vast majority of abortions are performed at clinics (as opposed to a doctor's office or hospital), so if more mothers are considering abortion, their contact with clinics increases.[38] Some who eventually decide against abortion may simply begin their pre-natal care at, say, the same Planned Parenthood clinic where they had received information about abortion.

[37] In fact, in 2012 Governor Rick Perry made a failed attempt to impose such a restriction. See `http://www.austinchronicle.com/blogs/news/2012-10-19/planned-parenthood-out-but-docs-can-still-discuss-abortion-in-new-texas-womens-health-program/`. While Perry failed in the attempt to restrict discussion of abortion, he succeeded in excluding Planned Parenthood from receiving funding through a Medicaid-waiver program.

[38] See Finer and Henshaw (2003). They report that 93 percent of abortions in 2000 were performed at clinics.

6.3 Do plans actively discourage high-cost births?

While it is in plans' interest to limit high-cost births, it could also be the case that the "missing" black infants are merely the result of MMC being better at providing access to contraception than FFS, and black mothers having had the greatest unmet demand under FFS. Though it is worth noting that Kearney and Levine (2009) find, nationally, when low-income women are provided greater access to contraception, Hispanic births fall differentially, suggesting Hispanics, not blacks, have greater unmet demand for contraception. Moreover, the increase in black abortions suggest if anything a rise in unwanted black conceptions under MMC, inconsistent with greater access to contraception.

In addition, it is perhaps harder to argue that, within the groups most likely to be on Medicaid, higher-cost subgroups such as older mothers differentially had the greatest unmet demand for contraception under FFS. Decreases in births among high-cost subgroups might be seen instead as evidence of a more active role played by plans. Indeed, recall that in Table 5 we saw that the share of births to mothers over 35 fall significantly for both blacks and Hispanics (by eight and four percent, respectively). There is no such pattern for our main placebo groups: foreign-born blacks and Hispanics and married whites (see col. 7 of, respectively, Appendix Tables 4, 5 and 6).

In summary, after MMC, unprofitable births become rarer. Black births—nearly twice as costly as Hispanic births—fall significantly after MMC. And among blacks and Hispanics, births to older mothers also fall significantly. We suspect these results arise from some mix of high-cost mothers' limiting fertility in response to poor care and outcomes and plans differentially discouraging their births, but we do not have the ability to separate these two mechanisms.

Comparing our results with past work on California MMC nicely illustrates the classic trade-off between incentives for cost-control and risk-selection. As noted, California has lower-powered incentives, as expensive cases were "carved out" and passed back to the state. Aizer *et al.* (2007) credit the carve-out with reducing incentives for preventive care, thus leading to worse birth outcomes under MMC. As the state both paid capitation payments to plans and picked up the bill for some adverse outcomes, total costs went up (Duggan, 2004). However, neither paper finds evidence that high-cost births went "missing" after MMC or that health disparities widened, as we do. Thus, carving out high-cost births appears to diminish plans' incentive to reduce long-run costs through preventive care, but it also decreases the incentive to avoid or discourage high-cost births, exactly the tradeoff described in Newhouse (1996).

7 Why might MMC increase care and health disparities between high- and low-cost groups? A simple theoretical model

Our empirical results presented above consistently document that the switch from FFS to MMC led to an increase in the health and care disparities between black and Hispanic pregnant women and infants. While our data do not allow us to precisely observe the mechanisms by which this effect might occur, we argue that a simple model of insurer incentives under MMC versus FFS generates predictions that are consistent with our findings.

7.1 Modeling incentives under MMC versus FFS

Consider two types of patients, healthy (H) and sick (S). Patient types are fixed over time. There are two types of costs that plans incur: those associated with preventive care θ and those associated with outcomes $c_i(\theta)$, where c varies by patient type.[39] For simplicity, let $c_H(\theta) = c(\theta)$ and $c_S(\theta) = c(\theta) + \alpha$, with $c' < 0$ and $c'' > 0$, so the returns to preventive care are the same across patient type. We do not distinguish between mothers and children and combine costs for both (empirically, there is very limited variation for costs related to the mother, and almost all variation comes from costs related to the infant).

Incentives under MMC. Under MMC, there are at least two plans from which patients can choose. Plans receive a capitation payment p regardless of patient type. Plans face a dynamic problem—how they treat a patient today determines whether she will return in the

[39] As both these costs are direct functions of θ we could instead formulate the model in terms of a total cost function, but splitting costs in this manner aids with intuition and maps more closely to the empirical results.

next period. In our context, "returning the next period" can either mean that the mother continues using this plan for the infant's later health care needs or that she returns to this plan the next time she is pregnant (and thus eligible for Medicaid herself). Let $\lambda(\theta)$ be the probability a patient choses the same plan in the next period, which is increasing concavely in the care she receives in the current period, so $\lambda' > 0$ and $\lambda'' < 0$. We scale down this probability by a discount factor δ to reflect the fact that she may exit the Medicaid program (e.g., no longer meet the income test) and to ensure a finite stream of expected profits.

We assume that plans can quickly learn patient type after a mother enrolls. First, they might form a reasonable estimate based on basic observables such as age and race. Second, in an initial check-up, information such as BMI, blood pressure, and health history will be gained. Third, diagnostic procedures throughout the pregnancy may reveal even more detailed information. We thus assume that patient type is observable to the plan at the point they are making many of their decisions about approving pre- and post-natal care.

Knowing patient type, each plan solves the following dynamic maximization problems:

$$V_t^H = \max_\theta \left\{ p - \theta - c(\theta) + \delta\lambda(\theta)V_{t+1}^H \right\} \quad [Healthy]$$

$$V_t^S = \max_\theta \left\{ p - \theta - c(\theta) - \alpha + \delta\lambda(\theta)V_{t+1}^S \right\} \quad [Sick]$$

Because for all θ, $p - \theta - c(\theta)$, the flow payoff from covering type H, is greater than $p - \theta - c(\theta) - \alpha$, the flow payoffs of covering type S, it holds that $V_{t+1}^H > V_{t+1}^S$. Differentiating each of the above expressions with respect to θ yields the following first-order conditions:

$$1 = -c'(\theta) + \delta\lambda'(\theta)V_{t+1}^H \quad [Healthy]$$

$$1 = -c'(\theta) + \delta\lambda'(\theta)V_{t+1}^S \quad [Sick].$$

For healthy patients, plans equate the marginal cost of an additional unit of θ (one) against two marginal benefits: that increasing θ decreases outcome costs—i.e., $-c'(\theta)$—while increasing the probability that the plan will enjoy the expected future profit stream—i.e., $\delta\lambda'(\theta)V_{t+1}^H$.

For sick patients, the incentives are the same, except that the continuation payoff $\delta\lambda'(\theta)V_{t+1}^{S}$ is smaller than that associated with a healthy patient, or perhaps negative. Either way, $V_{t+1}^{H} > V_{t+1}^{S}$ and $c'' > 0$ and $\lambda'' < 0$, so it must be that $\theta_H^{MMC*} > \theta_S^{MMC*}$.

Incentives under FFS. For simplicity, we model providers under FFS as being completely indifferent to outcome costs c_i—they merely send the bills back to the state. We assume that FFS providers get paid some reimbursement rate ρ for θ, and their cost of effort (or opportunity cost) is $e(\theta)$, which is increasing convexly in θ. Thus, for each client they provide some standard amount of care that satisfies $\rho = e'(\theta)$, and so $\theta_H^{FFS*} = \theta_S^{FFS*} \equiv \theta^{FFS*}$.

Predictions. The key result of the model is a divergence of health inputs θ for healthy and sick groups under MMC relative to FFS. That is:

$$(\theta_H^{MMC*} - \theta_H^{FFS*}) > (\theta_S^{MMC*} - \theta_S^{FFS*}).$$

Assuming that health inputs have the expected effect on health outcomes, we predict the same divergence in outcomes after the switch from FFS to MMC—outcomes for healthy clients improve while those for sick clients deteriorate.

As discussed, managed care plans under capitation have an incentive to compete for low-cost enrollees and appear unattractive to high-cost enrollees in the hope they will join a competing plan in the next period. For example, in our context, MMC plans would have an incentive to provide discretionary in-home visits or free transportion (some of the "voluntary benefits" mentioned in the Texas Medicaid documentation) to clients whose future pregnancies have expected costs below the capitation payment, so as to retain them as clients. By contrast, they would want high-cost clients to sign up with a competing plan during their next pregnancy and would thus ration such care. Moreover, they may make it more difficult for high-cost groups to enroll (e.g., demanding more detailed documentation proving eligibility). As such, this framework can explain why outcomes for high-cost groups can deteriorate relative to low-cost groups, consistent with what we find in the data.

Additionally, this model implicitly predicts that the effective price of childbearing increases for high-cost groups, while decreasing for low-cost groups. As such, the switch to MMC can affect birth composition as the groups whose care diminishes under MMC may lower their fertility (either through lower conception rates or higher abortion rates) in response. Albanesi and Olivetti (2010) offer evidence that improved health care for pregnant women during the 1950s contributed to the Baby Boom.[40] Moreover, if the continuation probability λ is not very responsive to quality of care θ and thus mothers' inertia is high, then plans might differentially encourage birth control (which is covered under Medicaid) for high-cost mothers.

Caveats to the model. Some of the simplifications and assumptions of the above framework deserve further discussion.

First, we do not model plans' incentives to influence initial enrollment—mothers find themselves in a certain plan and then make decisions about future enrollment based on the care they received in the plan. In fact, as in the model of Glazer and McGuire (2000), plans may design their benefits to deter sick individuals from enrolling *ex-ante*. Assuming that they do, it seems likely that they would still engage in the *ex-post* risk-selection activities we model after patient type is further revealed. Moreover, as we have documented, race and ethnicity are critical determinants of cost, so that how plans treat a high- (low-) cost patient will likely feedback to who enrolls initially. For example, improving the care of a Hispanic mother will increase the chances she recommends the plan to her friends and family (who are also very likely to be Hispanic) and past work has documented that individuals rely heavily on such recommendations in choosing a health plan.[41]

Second, we assume that utilization of care, health outcomes, and client retention all

[40] There is a small literature on whether Medicaid itself or similar programs that provide pre- or post-natal care are pro-natalist. As discussed by Lopoo and Raissian (2012), as Medicaid has generally provided both enhanced coverage for the costs related to child birth *as well as* access to birth control, it is hard to separate whether the enhanced coverage alone would be pro-natalist.

[41] Edgman-Levitan and Cleary (1996) document that seniors value word-of-mouth recommendations from friends and family more than they do aggregate "report card"-type ratings in choosing a managed care plan. Isaacs (1996) surveys adults of all ages and finds that family and friends' recommendations are weighed nearly the same as a *doctor's* recommendation in choosing a plan.

positively covary. It could instead be the case that utilization does not influence outcomes (e.g., if we are at the "flat of the curve," as in Fuchs, 2004), though, as previously noted, Aizer *et al.* (2007) found that MMC-induced reductions of care were large enough to affect mortality and other outcomes. Similarly, it could be that the aspects of health care that induce patient satisfaction and retention are not those that influence health outcomes.[42] Our model implicitly assumes there is sufficient overlap between the aspects of care that positively influence birth outcomes and those that lead mothers to continue with the MMC plan.

Third, to replicate the MMC setting, consumers pay zero premiums in our model. Future work might generalize this framework to allow consumers to pay more for enhanced coverage, as in the ACA exchanges. Depending on the correlations between income, insurance demand, and health, we suspect that allowing variation in premiums can either increase or decrease risk-selection.

Finally, our model does not give predictions on average care or outcomes under FFS versus MMC. For example, if ρ (the FFS reimbursement rate) is very low, then the level of care θ might increase for both high- and low-cost groups under MMC. While Aizer *et al.* (2007) found deleterious effects on average quality of care in California, predictions on average outcomes are outside our model. As such, welfare effects are not straightforward, an issue we discuss further in the next section.

8 Conclusion

We examine the experiences of black and Hispanic pregnant women and infants—two groups that have observably large differences in average healthcare costs and who are disproportionally covered by public health insurance—in an "exchange" setting where the government finances and regulates competing capitated private insurance plans but does not itself ad-

[42] An interesting example of an aspect of care that generates patient satisfaction but has no clear effects on outcomes is whether the race of the doctor and patient match. Past work on doctor-patient "race-concordance" finds that it strongly influence patient satisfaction (LaVeist and Nuru-Jeter, 2002), modestly influences health care utilization (LaVeist *et al.*, 2003) but has no clear relationship with patient outcomes (Meghani *et al.*, 2009).

minister a FFS plan. We focus on the transition from FFS Medicaid to Medicaid managed care in Texas to measure the causal effects of MMC on care provision and health outcomes among black and Hispanic births. Our results show that the black-Hispanic mortality, low birth weight and pre-term birth gaps increase by 42, 13 and 22 percent, respectively, after a county switches from FFS to MMC. Quality of pre-natal care and birth procedures generally improve for Hispanics relative to blacks and black birth rates fall substantially after MMC.

We offer an explanation of our empirical findings through a simple dynamic model of risk-selection in these types of "exchange" settings. In our model, plans have incentives to retain healthy, low-cost patients, whereas they prefer their high-cost clients to switch to a competing insurer. As such, they improve care for the former group at the expense of the latter.

While our results provide strong evidence for how care and outcomes diverge for high- and low-cost groups under MMC, they do not directly speak to overall welfare. Given the larger number of Hispanics in Texas, average birth outcomes do not decline despite significant deterioration among black infants. However, if society wishes to shrink health disparities, then MMC may be inferior to FFS as it transfers health resources away from a group with poor average health. As the returns to health investments are thought to be lower for the healthy than the sick (Grossman, 1972), such a transfer could lower total welfare. Finally, if health insurance is in part meant to smooth the utility consequences of *ex-ante* differences in health, then, relative to FFS, MMC's transferring of resources from the sick to the healthy weakens the insurance value of the Medicaid program.

The welfare effects of changes in birth composition are even more difficult to interpret. Given the challenges single-parent households face, that the decline in black births is driven by unmarried mothers could be seen as a positive effect of the reform. However, this view may be too narrow and the desirability of the result may depend on the mechanism—e.g., whether plans are merely supplying birth control to women with previously unmet demand or if they are actively discouraging births among mothers because of their expected costs.

Future work can hopefully make more progress toward pinning down underlying mechanisms than we have. In fact, this limitation is generally shared by the literature on risk-selection: there is ample evidence that private capitated plans often manage to avoid high-cost enrollees, but almost no evidence on *how* they accomplish this selection. To our knowledge, there is only one audit study on risk-selection, Bauhoff (2012). He finds that even highly regulated private plans in the German health system are slower to enroll individuals who contact them from high-cost regions of the country. More work along these lines seems essential, especially in the U.S. context where the ACA will soon provide millions of individuals health coverage through private, capitated plans.

In our model, an inefficiency arises because plans want clients with costs above the capitation payment to switch to a competitor and thus reduce their care below the socially-optimal level. This result suggests that competition may undermine the underlying policy goal of capitation—instead of acting as the residual claimant on costs above or below the capitation payment and thus internalizing patients' future costs, plans attempt to pass on these costs to their competitor. This externality problem would not exist with a monopolistic insurer (though other problems associated with monopoly might arise). As all counties in Texas offer a choice between at least three competing MMC plans, we could not compare counties with and without competition, but future work might examine other MMC settings.

Introducing risk-adjustment could potentially address the risk-selection results we have presented, though, historically, governments have been reluctant to risk-adjust based on race. Adjusting based on past health conditions is very challenging in the MMC setting and, as previously noted, is rarely attempted. First, plans would have to submit some accounting of their clients' health conditions to the government, so "intensive" coding becomes a problem, as it is in Medicare Advantage and the Medicare Prospective Payment System.[43] Second, while Medicare can calibrate a risk-adjustment formula by regressing enrollee costs

[43] See the Center for Medicare and Medicaid Services on "intensive coding" among MA plans: `http://www.cms.gov/Medicare/Health-Plans/MedicareAdvtgSpecRateStats/downloads/Advance2008.pdf`. See Silverman and Skinner (2004) on provider "up-coding" practices after PPS.

on dummies for past health conditions using cost and claims data from its FFS pool, state governments under MMC typically do not have a public FFS option and thus will not have the cost and claims data that Medicare uses. Third, risk-adjustment formulae typically document existing health conditions using twelve months of pre-data, and use this information to forecast costs for the following twelve months. A stable client population—as in Medicare—is thus required, a challenge in Medicaid given the "churning" of the client base.

Each of these three challenges would seem to apply equally to the ACA state exchanges. Regulators will rely on insurers' accounting of client health conditions, and as there is no public FFS option they will not have their own cost and claims data. And as the exchanges serve those too rich for Medicaid but not so well off as to have employer insurance, their clients will likely come and go based on outside options—if their situation improves, then they will move into employer insurance; if their situation deteriorates, then they may need to switch to Medicaid.

With Medicaid Managed Care, the ACA exchanges, Medicare Part D and the prominence of Medicare premium-support proposals, the U.S. is moving rapidly toward providing public health insurance through a model of competing, capitated private insurance plans. Past work has identified challenges associated with this model, including the increase in costs that come with insurers losing monopsony bargaining power over providers and consumers' cognitive overload from choosing among a large set of options.[44] Our work points to an additional concern arising from consumer choice—it tempts insurers to under-serve high-cost clients in the hope they will switch to a competitor. However, in most contexts consumer choice and competition are beneficial, and restricting choice among insurers all else equal has been found to significantly decrease consumer surplus.[45] Given the direction of U.S. health policy, future work to better assess these trade-offs is of growing importance.

[44] See Dafny *et al.* (2012) and Abaluck and Gruber (2011), respectively.
[45] See Dafny *et al.* (2013).

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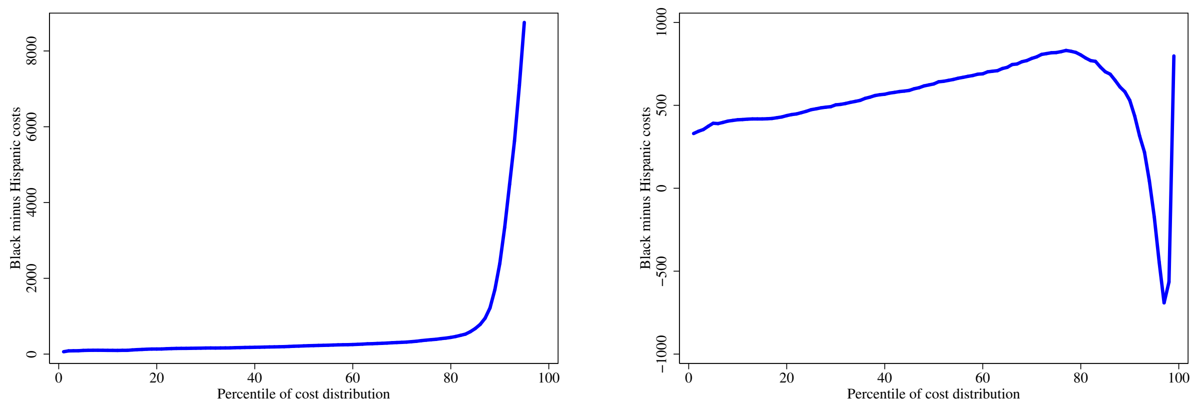
Figure 1: Distribution of hospital charge differences between Blacks and Hispanics

(a) Newborns

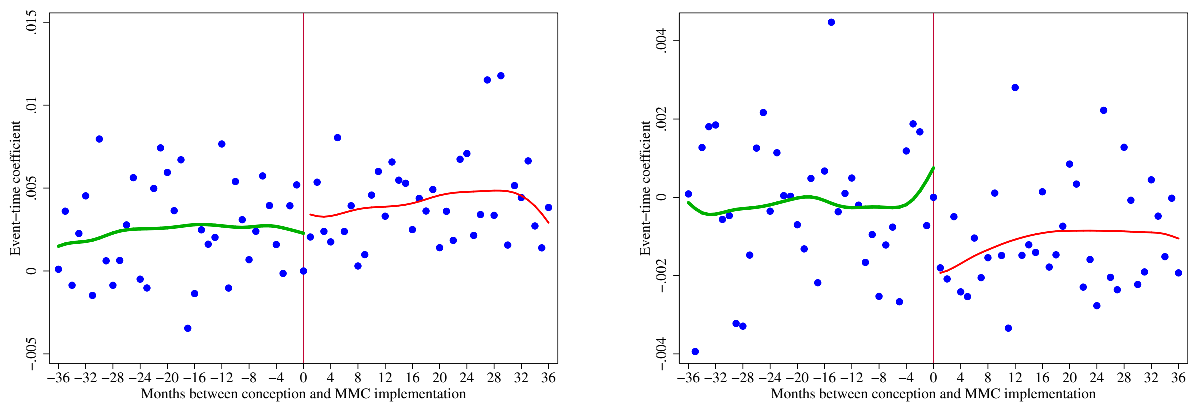
(b) Deliveries



Notes: Figures are based on data from public-use Texas Hospital discharge data (see `http://www.dshs.state.tx.us/THCIC/Hospitals/Download.shtm` to download these data). For each graph, the value of the Hispanic n^{th} percentile is subtracted from the value of the Black n^{th} percentile. Because of the extreme skewness of the newborn charges, the graph is truncated at the 95^{th} percentile. The black-Hispanic difference for the 99^{th} percentile is \$76,341.

Figure 2: Changes in mortality to infants born to U.S.-born black and Hispanic mothers (note different scales)

(a) Blacks (b) Hispanics



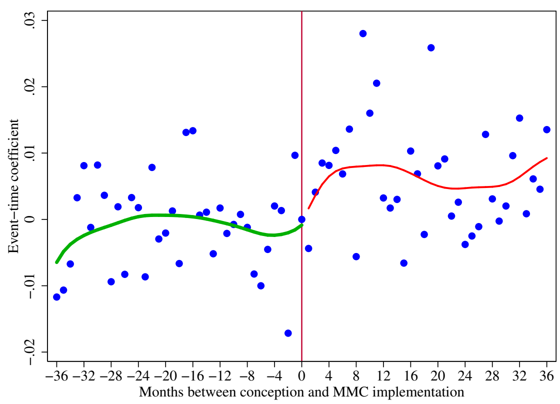
Notes: These figures show the results from estimating mortality rates for black (Figure a) and Hispanic (Figure b) births in the 36 months before and after MMC implementation. Specifically, we estimate the following equation:

$$Y_{ymc} = \sum_{n=-36}^{-1} \beta_n \mathbf{I}_{ct}^n + \sum_{n=1}^{36} \beta_n \mathbf{I}_{ct}^n + \eta Window_{ct} + \mu_c + \gamma_y + \nu_m + \mu_c * t + \epsilon_{ymc},$$

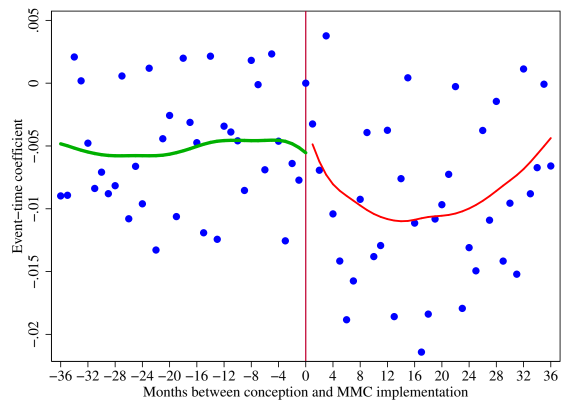
where $\mathbf{I}_{tc}^n$ is an indicator variable for conceptions n months after a county c switched to MMC, meaning negative values of n indicate conceptions in months *before* MMC implementation. $Window$ is an indicator for being conceived within a six-year window of MMC's introduction (the range of the figure). This addition allows us to normalize conceptions the same month as MMC implementation to zero for ease of interpretation. (Excluding $Window$ only shifts the level, not the shape, of the figures, as the excluded group by default becomes all births outside the graphs' six-year window.) The figure plots the β_n coefficients along with lowess lines (of bandwidth one). Otherwise, the notation follows exactly from our main estimating equation (1) in the text: c indexes counties, and y and m month and year; Y_{ymc} is an outcome measure for county c in year-month y-m (in this case, black and Hispanic mortality, respectively); μ_c are county fixed effects, γ_y are conception-year fixed effects, and λ_m are conception-month fixed effects; $\mu_c * t$ is the county-specific linear time-trend.

Figure 3: Changes in pre-term share of births to U.S-born black and Hispanic mothers (note different scales)

(a) Blacks



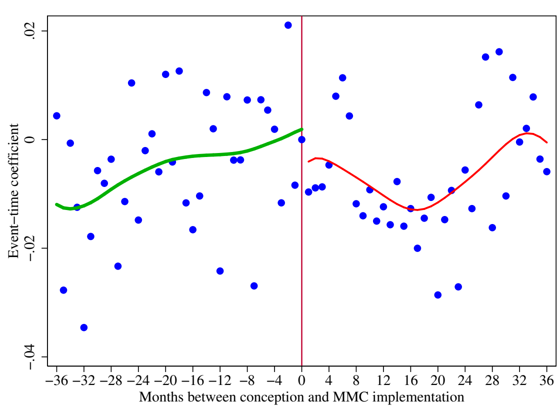
(b) Hispanics



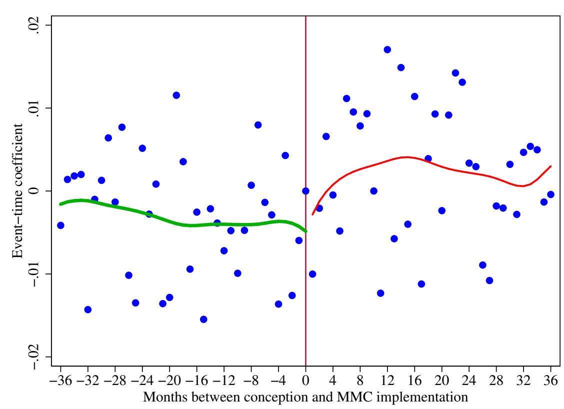
Notes: These figures are identical to those displayed in Figure 2 except that pre-term birth serves as the outcome variable.

Figure 4: Changes in the male share of births to U.S-born black and Hispanic mothers (note different scales)

(a) Blacks

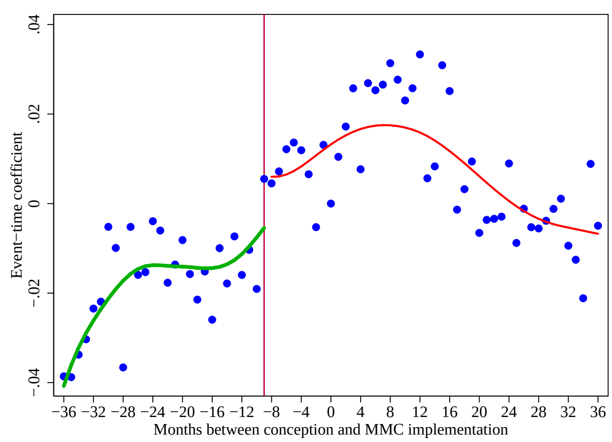


(b) Hispanics



Notes: These figures are identical to those displayed in Figure 2 except that share male serves as the outcome variable.

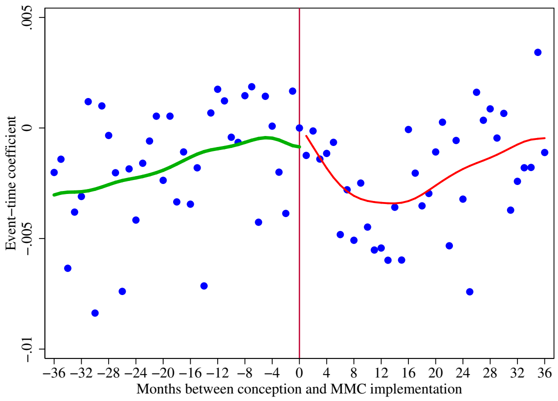
Figure 5: Share receiving pre-natal care in a public clinic, U.S.-born black mothers



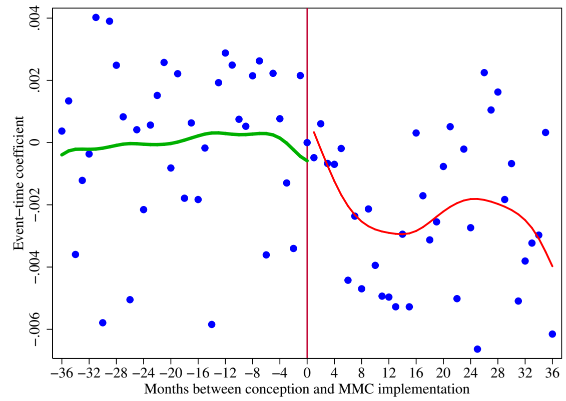
Notes: This figures are identical to those displayed in Figure 2 except that share receiving pre-natal care in a public clinic serves as the outcome variable and the pre- and post-period are defined as births (not conceptions) before and after MMC.

Figure 6: Black share of all births before and after MMC

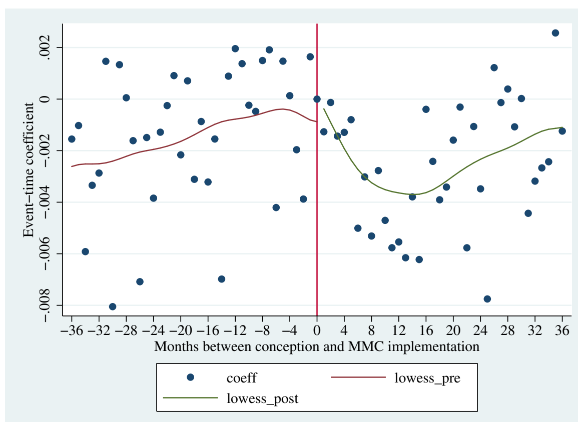
(a) With country linear time trends



(b) Without county time trends



(c) With county quadratic time trends



Notes: These figures show the results from estimating the effects on the share of black births in the 36 months before and after MMC implementation (the month of MMC implementation is normalized to zero). See the notes to Figure 2 for further details on the estimation procedure. Figure (a) includes country linear time trends (our standard specification), Figure (b) includes no county time trends, and Figure (c) includes county quadratic time trends.

Table 1: Hospital charges for newborns and deliveries

	Newborn	Delivery			
	(1)	(2)	(3)	(4)	(5)
Black	4218.3***	1485.8***	1499.7***		
	[110.0]	[16.52]	[16.51]		
Age 35 or older				1130.5***	963.4***
				[60.70]	[22.89]
Mean, dept. var.	5813.6	7107.5	7107.5	7608.6	7002.8
Mean, ex. group	5236.6	7002.9	7002.9	7510.8	6916.9
Pct. diff	0.806	0.212	0.214	0.151	0.139
Age cat. FE	No	No	Yes	No	No
Sample	Bl., H.	Bl., H.	Bl., H	Bl.	H.
Observations	816914	788637	788637	144403	645682

Notes: Regressions are based on data from public-use Texas Hospital discharge data (see `http://www.dshs.state.tx.us/THCIC/Hospitals/Download.shtm` to download these data). All regressions include county and year fixed effects and include all Hispanic and black births from the third quarter of 1999 through 2004 (county identifiers are missing in the first two quarters of 1999). Col. (3) includes maternal age fixed effects ($age < 20, age \in [20, 25), age \in [25, 30), age \in [30, 35), age \geq 35$). All means of the dependent variable are reported, as well as the percent difference between the group denoted by the reported regression coefficient (e.g., blacks, in col. 1) and the excluded group (e.g., Hispanics, in col. 1). That is, "Pct. Diff" just divides the coefficient by the excluded-group mean. Cols. (1) through (3) include all blacks and Hispanics, col. (4) includes only blacks and col. (5) includes only Hispanics.

Table 2: Summary statistics

	(1) All	(2) U.S. Bl.	(3) U.S. Hisp.	(4) For. Bl.	(5) For. H.	(6) Mar. Wh.
Mother's age	25.76	24.12	23.79	28.72	25.93	28.05
	(6.063)	(5.949)	(5.835)	(5.902)	(5.789)	(5.555)
Child died (death cert. matched to birth cert.)	0.00725	0.0120	0.00715	0.0135	0.00565	0.00614
	(0.0848)	(0.109)	(0.0843)	(0.116)	(0.0750)	(0.0781)
Pre-term (Gestation less than 37 weeks)	0.0923	0.135	0.0959	0.114	0.0755	0.0859
	(0.289)	(0.342)	(0.294)	(0.318)	(0.264)	(0.280)
Low-birth weight (Birthweight below 2,500 g.)	0.0724	0.127	0.0733	0.0983	0.0579	0.0599
	(0.259)	(0.333)	(0.261)	(0.298)	(0.234)	(0.237)
Birthweight below 2,500 g. or above 4,000 g.	0.159	0.172	0.142	0.192	0.148	0.174
	(0.365)	(0.378)	(0.349)	(0.394)	(0.355)	(0.379)
Male	0.511	0.509	0.510	0.505	0.510	0.513
	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)
Pre-natal care in first month	0.229	0.210	0.219	0.248	0.171	0.293
	(0.420)	(0.407)	(0.414)	(0.432)	(0.376)	(0.455)
Pre-natal care at public clinic	0.126	0.136	0.124	0.0983	0.262	0.0398
	(0.332)	(0.343)	(0.330)	(0.298)	(0.440)	(0.195)
Pre-natal care at hospital	0.172	0.248	0.134	0.292	0.294	0.0959
	(0.378)	(0.432)	(0.341)	(0.455)	(0.456)	(0.294)
Pre-natal care at private doctor's office	0.677	0.601	0.745	0.581	0.362	0.851
	(0.467)	(0.490)	(0.436)	(0.493)	(0.481)	(0.356)
Observations	2,814,681	296,589	646,053	21,555	617,608	922,142

Notes: This table reports means for key variables in the Texas birth records data. The sample of analysis includes births that were conceived by mothers residing in Texas between January 1993 and December 2001.

Table 3: Effect of MMC on mortality rates ($\times 100$) for U.S.-born black and Hispanic births

	(1) Bl.	(2) Hsp.	(3) Bl.	(4) Hsp.	(5) Bl.	(6) Hsp.
Conceived after MMC	0.139**	-0.154**	0.142**	-0.149**	0.260**	-0.169**
	[0.0647]	[0.0722]	[0.0639]	[0.0650]	[0.112]	[0.0675]
Log county pop.			1.142	-5.976***	-3.955	-5.368*
			[3.466]	[1.983]	[6.172]	[2.838]
Log per cap. county income			2.709	-1.389**	5.437**	-0.553
			[1.898]	[0.632]	[2.501]	[1.197]
County unemp. rate			1.196	-0.610	-1.673	-0.361
			[5.375]	[1.275]	[5.801]	[2.180]
Dept. var mean	1.198	0.715	1.198	0.715	1.260	0.822
Sample	All	All	All	All	Unmar.	Unmar.
Diff/p-val	0.293	0.00110	0.291	0.000708	0.429	0.00281
Reg. obs (cells)	12833	20504	12833	20504	11766	16370
Indiv. obs.	296589	646053	296589	646053	190899	250154

Notes: These regressions are based on Texas birth records data. The sample of analysis includes births that were conceived by mothers residing in Texas between January 1993 and December 2001. Units of observation are county/conception-year/conception-month cells and all regressions are weighted by cell size. All regressions include year, month and county fixed effects, and county-specific linear time trends. Standard errors are clustered by county. The "Diff/p-val" row shows in the odd-numbered columns the differences in the black-Hispanic MMC coefficients and the even-numbered columns present the p-value associated with the test of equality across the two coefficients.

Table 4: Effect of MMC on other birth outcomes (×100) for U.S.-born black and Hispanic births

	Preterm		LBW		Abn. BW		Male	
	(1) Bl.	(2) Hsp.	(3) Bl.	(4) Hsp.	(5) Bl.	(6) Hsp.	(7) Bl.	(8) Hsp.
Conceived after MMC	1.013***	-0.607***	0.705*	0.0231	0.906***	-0.146	-0.627	0.779***
	[0.342]	[0.175]	[0.376]	[0.175]	[0.330]	[0.363]	[0.412]	[0.245]
Dept. var mean	13.51	9.593	12.72	7.334	17.25	14.21	50.95	51.04
Diff/p-val	1.621	1.23e-10	0.682	0.0254	1.052	0.0165	-1.406	0.00335
Reg. obs (cells)	12833	20504	12828	20502	12828	20502	12833	20504
Indiv. obs.	296589	646053	296584	646051	296584	646051	296589	646053

Notes: See notes under Table 3 for more details about the data, sample, and specifications. "LBW" denotes birth weight $< 2,500$g; "Abn. BW" (abnormal birthweight) denotes birthweight $< 2,500$ g or >4,000 g; "Pre-term" denotes gestation < 37 weeks; "Complic" denotes any birth complications.

Table 5: Changes in risk-factors (×100) after MMC for U.S.-born black and Hispanic mothers

	Older		Diab/Hyper.		Smokes		Pre. mort. I		Pre. mort. II	
	(1) Bl.	(2) Hsp.	(3) Bl.	(4) Hsp.	(5) Bl.	(6) Hsp.	(7) Bl.	(8) Hsp.	(9) Bl.	(10) Hsp.
Conceived after MMC	-0.552**	-0.349**	-0.111	-0.247	-0.335	0.267	-0.00972***	-0.00105	-0.0110***	-0.00220
	[0.238]	[0.145]	[0.191]	[0.216]	[0.254]	[0.211]	[0.00288]	[0.00134]	[0.00367]	[0.00142]
Dept. var mean	5.659	4.699	3.469	3.164	6.284	3.483	1.182	0.708	1.176	0.706
Diff/p-val	-0.203	0.398	0.136	0.593	-0.602	0.0196	-0.00867	0.00607	-0.00881	0.0123
Reg. obs (cells)	12832	20504	12833	20504	12808	20489	12801	20471	12776	20456
Indiv. obs.	296588	646053	296589	646053	296563	646037	296556	646019	296530	646003

Notes: See notes under Table 3 for more details about the data, sample, and specifications.

Table 6: Effect of MMC on pre-natal care measures ($\times$100) for U.S.-born black and Hispanic births

	Imm. PNC		PVS		PVS>7		Δ W >15		Δ W >20	
	(1) Bl.	(2) Hsp.	(3) Bl.	(4) Hsp.	(5) Bl.	(6) Hsp.	(7) Bl.	(8) Hsp.	(9) Bl.	(10) Hsp.
Born under MMC	-2.000	0.0384	-0.0817	-0.0642	-2.261**	-0.850	-1.276	0.0808	-2.013**	0.163
	[1.730]	[0.852]	[0.0734]	[0.0697]	[1.033]	[0.733]	[0.851]	[0.689]	[0.813]	[1.108]
Dept. var mean	21.01	21.88	10.45	10.87	79.42	83.02	86.45	87.11	74.48	74.72
Diff/p-val	-2.039	0.0547	-0.0176	0.812	-1.410	0.0272	-1.357	0.0845	-2.176	0.101
Reg. obs (cells)	12767	20424	12617	20271	12617	20271	12192	19902	12192	19902
Indiv. obs.	296516	645966	296225	645741	296225	645741	295429	645237	295429	645237

Notes: See notes under Table 3 for more details about the data, sample, and specifications. Note that the key explanatory variable of interest is an indicator for being born after (rather than conceived after) MMC. "Imm. PNC" denotes "immediate pre-natal care," indicating that the mother received care within the first month of her pregnancy. "PVS" denotes the total number of pre-natal care visits. "PVS>7" denotes more than 7 visits. The remaining two outcomes refer to maternal weight gain (ΔW) in pounds.

Table 7: Effect of MMC on site of pre-natal care ($\times$100) for U.S.-born black and Hispanic births

	Pub. Clinic		Hosp.		Private	
	(1) Bl.	(2) Hsp.	(3) Bl.	(4) Hsp.	(5) Bl.	(6) Hsp.
Born under MMC	3.573**	1.198	-1.495**	0.511	-0.620	-1.269
	[1.621]	[1.348]	[0.610]	[1.366]	[0.705]	[1.559]
Dept. var mean	13.64	12.43	24.88	13.42	59.98	74.43
Diff/p-val	2.375	0.0238	-2.006	0.122	0.648	0.638
Reg. obs (cells)	12535	20039	12535	20039	12535	20039
Indiv. obs.	290795	635297	290795	635297	290795	635297

Notes: See notes under Table 3 for more details about the data, sample, and specifications. Note that the key explanatory variable of interest is an indicator for being born after (rather than conceived after) MMC. "Pub. Clinic" denotes pre-natal care received at a public clinic, "Hospital" denotes pre-natal care received at a hospital, and "Private" denotes pre-natal care received at a private doctor's office.

Table 8: Effect of MMC on U.S.-born black birth rates

	Share of births		Log of births		
	(1) Black	(2) Bl. unm.	(3) Black	(4) Bl. unm.	(5) Non-bl.
Conceived after MMC	-0.00120	-0.00247***	-0.0309**	-0.0462***	-0.0141
	[0.000790]	[0.000630]	[0.0130]	[0.00843]	[0.00920]
Mean, dept. var.	0.105	0.0678	5.526	5.091	7.070
Reg. obs. (cells)	26021	26021	3672	3672	3672
Indiv. obs.	2814681	2814681	258480	164943	1638601

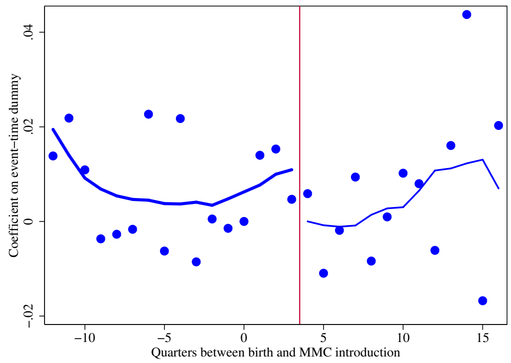
Notes: These regressions are based on Texas birth records data. The sample of analysis includes births that were conceived by mothers residing in Texas between January 1993 and December 2001. Units of observation are county/conception-year/conception-month cells. In all columns except when log of births is the outcome, all county/year/month cells in Texas between 1993 and 2001 are included and each cell is weighted by cell size. When logs are used in cols. (4) through (6), counties are restricted to those with at least one black unmarried birth in each month (to avoid taking the log of zero and to have a consistent sample of counties), a sample which accounts for 67 percent of all births and 87 percent of black unmarried births. Col. (4) is weighted by the number of black births in a county/year/month, col. (5) is weighted by the number of black unmarried births in a county/year/month, and column (6) is weighted by the number of non-black births in a county/year/month. All regressions include year, month and county fixed effects, and county-specific linear time trends. Standard errors are clustered by county.

Table 9: Effect of MMC on abortion rates (×100)

	Dept. var: Share of conceptions ending in abortion							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Black	Bl. unm.	Bl. mar.	Bl. unm.	Bl. unm.	Hisp	H. unm.	W. mar
Conceived after or 3	1.886**	3.083***	-0.682	2.809***	1.404**	-0.242	0.258	0.00430
mos. before MMC	[0.795]	[0.752]	[0.636]	[0.721]	[0.607]	[0.600]	[1.231]	[0.00369]
Mean, dept. var.	27.87	34.67	13.29	34.68	34.67	12.79	24.77	0.0520
County trends	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes
Early cut-off	No	No	No	Yes	No	No	No	No
Reg obs. (cells)	6134	5649	4076	5406	5649	10397	8913	8222
Indiv. obs.	198998	135732	63266	129950	135732	704902	278956	275522

Notes: These regressions are based on Texas abortion and birth records data. The sample of analysis includes births that were conceived by mothers residing in Texas between January 1998 and December 2001, since abortions only began to be reported then. Units of observation are county/conception-year/conception-month cells and all regressions are weighted by cell size. All regressions include year, month and county fixed effects and all but col. (5) include linear county time trends. Standard errors are clustered by county.

Appendix Figure 1: Black share of population ages 2-17 (IPUMS data)



Notes: These data are taken from the 2005-2011 IPUMS, restricted to individuals born in Texas. Linear birth-date (measured at the quarter level) county trends are included (the analogue to county linear trends in conception month in the birth-certificate analysis), as well as county and year- and quarter-of-birth fixed effects. Conception corresponds to births three-quarters after MMC implementation, so we divide the lowess lines at that point. The youngest cohort to be born during an MMC transition year would be five to six years old in 2005 (from those counties switching in 1999) and the oldest would be fourteen to fifteen in 2011 (from those counties switching in December 1995). As such, we include ages slightly below and above these cut-offs respectively, just as in the birth-certificate analysis we include about three years before and after the first and last set of counties switch, respectively.

Appendix Table 1: Roll Out Schedule for Texas MMC

Date	Counties
Aug 1993	Travis
Dec 1993	Chambers Jefferson Galveston
Dec 1995	Liberty, Hardin, Orange
Sep 1996	Burnet Williamson Lee Bastrop Fayette Caldwell
	Hays Lubbock Terry Lynn Garza Crosby Hockley Llano Hale
	Floyd Swisher Randall Deaf Smith Potter Hutshinson Carson
	Bexar Atascosa Wilson Guadalupe Comal Kendall Bandera
	Medina Tarrant Hood Parker Wise Denton Johnson
Dec 1997	Houston
Mar 1998	Harris Galveston Brazoria Matagorda Wharton Fort Bend
	Austin Waller Montgomery
Jan 1999	Dallas Ellis Navarro Kaufman Rockwall Hunt Collin El Paso
	Hudspeth
Jan 2006	Nueces Kenedy Brooks Kleberg Jim Wells San Patricio
	Live Oak Aransas Refugio Bee Goliad Victoria Karnes Calhoun

Notes: This information was obtained from Chapter 6 of the report available here: `www.hhsc.state.tx.us/medicaid/reports/PB8/PinkBookTOC.html`

Appendix Table 2: Is MMC rollout correlated with underlying county trends?

	(1) Log Pop.	(2) Log Per-Cap. Inc.	(3) Unemp. Rate
After MMC	0.00190	-0.00262	-0.0000603
	[0.00190]	[0.00357]	[0.00347]
Mean, dept. var	13.05	10.04	0.0613
Reg. obs. (cells)	26021	26021	26021
Underlying # births	2,814,681	2,814,681	2,814,681

Regressions include county and year fixed effects and county linear time trends.

Appendix Table 3: Estimated Medicaid share of births in 2005

	(1) All	(2) U.S. Bl.	(3) U.S. Hisp	(4) For. Bl.	(5) For. Hsp.	(6) Wh.
Medicaid share	0.539	0.836	0.877	0.338	0.271	0.437
Medicaid share, married	0.360	0.471	0.692	0.245	0.265	0.269
Observations	273,471	26,615	69,146	2,647	64,610	100,526

Notes: Texas does not record Medicaid status on birth certificates until 2005. As we discuss in Section 4, these numbers appear substantially under-reported, likely due to women or providers who are on privatized Medicaid mistakenly reporting that the birth is covered by a private or "other" instead of Medicaid. In 2005, the Texas Health and Human Services Commission reported that 54 percent of births were covered by Medicaid, whereas the birth certificate data indicate that only 41 percent were. We thus "gross up" the Medicaid share by 1.3 in this table. See Section 4 for additional evidence that the shift from FFS to MMC results in substantial under-reporting of Medicaid births.

Appendix Table 4: Effect of MMC on birth outcomes (×100) for foreign-born black mothers

	(1) Mort.	(2) Pret	(3) LBW	(4) ABW	(5) Male	(6) Older
Conceived after MMC	0.0305 [0.256]	-0.677 [0.849]	-0.869* [0.462]	0.124 [1.030]	2.667* [1.559]	0.522 [1.278]
Mean, dept. var	1.355	11.38	9.831	19.20	50.52	14.69
Reg. obs. (cells)	2387	2387	2381	2381	2387	2386
Underlying	21555	21555	21549	21549	21555	21554

Appendix Table 5: Effect of MMC on birth outcomes (×100) for foreign-born Hispanic mothers

	(1) Mort.	(2) Pret	(3) LBW	(4) ABW	(5) Male	(6) Older
Conceived after MMC	-0.0424 [0.0535]	0.234 [0.483]	-0.184 [0.181]	-0.238 [0.364]	-0.0727 [0.472]	0.000427 [0.133]
Mean, dept. var	0.565	7.550	5.794	14.78	51.01	7.459
Reg. obs. (cells)	18153	18153	18147	18147	18153	18152
Underlying	617608	617608	617602	617602	617608	617607

Appendix Table 6: Effect of MMC on birth outcomes (×100) for married white mothers

	(1) Mort.	(2) Pret	(3) LBW	(4) ABW	(5) Male	(6) Older
Conceived after MMC	0.0528 [0.0443]	0.261 [0.221]	0.0163 [0.156]	-0.0874 [0.186]	-0.115 [0.198]	0.0254 [0.187]
Mean, dept. var	0.614	8.589	5.991	17.44	51.27	11.45
Reg. obs. (cells)	23898	23898	23894	23894	23898	23898
Underlying	922142	922142	922138	922138	922142	922142

Appendix Table 7: Effect of MMC on U.S.-born black birth outcomes after controlling for covariates

	(1) Mort.	(2) Pret.	(3) LBW	(4) ABW	(5) Male
Conceived after MMC	0.146** [0.0654]	1.075*** [0.359]	0.872** [0.428]	1.042*** [0.358]	-0.611 [0.467]
Mean, dept. var.	1.198	13.51	12.72	17.25	50.95
Observations	296589	296589	296279	296279	296589

Notes: These regressions are based on individual-level Texas birth records data. The sample of analysis includes births that were conceived by mothers residing in Texas between January 1993 and December 2001. All regressions include the typical controls in the cell-aggregated regressions (county, year, and month fixed effects and county time trends) as well as the following individual-level controls: indicators for married and first-parity child, age (in four-year bins) fixed effects, and educational attainment fixed effects (no high school education, high school education, some college and college graduate). Standard errors are clustered by county.

Appendix Table 8: Effect of MMC on U.S.-born Hispanic birth outcomes after controlling for covariates

	(1) Mort.	(2) Pret.	(3) LBW	(4) ABW	(5) Male
Conceived after MMC	-0.150** [0.0716]	-0.554*** [0.168]	0.0415 [0.175]	-0.0857 [0.399]	0.710*** [0.235]
Mean, dept. var.	0.715	9.593	7.334	14.21	51.04
Observations	646053	646053	645778	645778	646053

Notes: These regressions are based on individual-level Texas birth records data. The sample of analysis includes births that were conceived by mothers residing in Texas between January 1993 and December 2001. All regressions include the typical controls in the cell-aggregated regressions (county, year, and month fixed effects and county time trends) as well as the following individual-level controls: indicators for married and first-parity child, age (in four-year bins) fixed effects, and educational attainment fixed effects (no high school education, high school education, some college and college graduate). Standard errors are clustered by county.

Appendix Table 9: Changes in black share of school enrollment after MMC

	(1) Share black	(2) Log bl. enrollment	(3) Log bl. enroll (w 0s)
After MMC	-0.000896	0.00362	0.00511
	[0.000871]	[0.0117]	[0.0124]
Mean, dept. var.	0.140	8.667	8.644
Number county-year cells	2738	2588	2738

Notes: These data come from the National Center of Education Statistics. Units of observation are county-year cells. All regressions are weighted by total enrollment in each cell. The sample of analysis includes school enrollment data from all Texas counties except for the four pilot counties over 1992-1993 to 2001-2002. In the "Log bl. enroll (w 0s)" specifications, cells with 0 values are recoded to 1. All regressions include county and year fixed effects and county-specific linear time trends. Standard errors are clustered by county.

Estimating the undocumented share of foreign-born Hispanic mothers in Texas

We calculate this share for the year 2000. According to the U.S. Census, there were $20,851,820$ residents in Texas in 2000.[46] According to the Pew Hispanic Center, there were 1.1 million undocumented immigrants in Texas in 2000.[47] Also according to Pew, 76 percent of undocumented immigrants nationwide are Hispanic, which is a vast underestimate for Texas, given its position on the U.S.-Mexican border.[48] As such, a lower bound for the number of undocumented Hispanics in Texas is $0.76 * 1,100,000 = 836,000$.

Using the 2000 IPUMS, we calculate that foreign-born Hispanics (regardless of their immigration status, which the Census does not record) account for 9.77 percent of the Texas population, or $0.0977 * 20,851,820 = 2,037,222$ people.

Finally, Pew notes that undocumented immigrants are 34 percent more likely to have children (the relevant group for our regression analysis) than are documented immigrants.[49] We thus gross up the estimated number of undocumented Hispanics in the first paragraph by 1.3.

Our final calculation of the share of Hispanic foreign-born mothers who are undocumented is thus $(1.3 * 836,000) \div 2,037,222 = 53.3$ percent. Again, because we assume that the Hispanic share of undocumented immigrants in Texas is equal to the national share, this calculation is a lower bound.

[46] See http://www.census.gov/population/www/cen2000/maps/files/tab02.pdf.

[47] See http://www.pewhispanic.org/2011/02/01/appendix-a-additional-figures-and-tables/.

[48] See http://www.pewhispanic.org/2009/04/14/a-portrait-of-unauthorized-immigrants-in-the-united-states/.

[49] See http://www.pewhispanic.org/2009/04/14/a-portrait-of-unauthorized-immigrants-in-the-united-states/.