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

Ilyana Kuziemko, Katherine Meckel and Maya Rossin-Slater

August 5, 2014

Abstract

Increasingly in U.S. public insurance programs, the state finances competing, capitated health plans rather than using a fee-for-service (FFS) model. We study how high- and low-cost infants (blacks and Hispanics, respectively) are affected by the transition from FFS to Medicaid managed care (MMC). We find that black-Hispanic infant health disparities *widen*—e.g., black mortality increases by 12% while the Hispanic mortality *decreases* by 22%—and care worsens for blacks. Additionally, black birth rates fall. We present a model of risk-selection in which capitation incentivizes competing plans to offer better care to low-cost clients to retain them in future periods.

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

---

*Princeton, Columbia and UC Santa Barbara Departments of Economics, respectively. Contact e-mail: kuziemko@princeton.edu. We are grateful to Janet Currie, Tal Gross, Jon Gruber, Kate Ho, Lisa Kahn, Melissa Kearney, Wojciech Kopczuk, Doug Miller and Andrea Prat for their feedback as well as seminar participants at Brookings, Columbia University, Northwestern, Princeton University, the University of Stavanger, the University of Virginia, Yale, the All-California Labor Economics Conference, the 3rd Annual Conference on Healthcare Markets, and the ASHEcon conference. We are indebted to Doneshia Ates, Marc Montrose, and Gene Willard at the Texas Department of State Health Services (DSHS) for assistance with the data, as well as Stephanie Goodman and David Palmer, the Spokesperson and Chief Actuary at the Texas DSHS, respectively. Abraham Bae and Jimmy Charite provided excellent research assistance. Meckel is grateful for support from a Graduate Research Fellowship from the National Science Foundation. The content is solely the responsibility of the authors and does not necessarily represent the official views of the National Science Foundation. All remaining errors are our own.

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) expands insurance coverage almost entirely through this new private model. The ACA insurance exchanges—the backbone of the reform—offer private, capitated, competing insurance plans with substantial government subsidies and regulation, but no public FFS option.

A key goal for any insurance program is to guarantee *ex-ante* high-cost patients appropriate levels of care, given the long recognized incentives for providers and insurers to cream-skim (Newhouse, 2006). In this paper, we seek to make three contributions in understanding whether "exchange settings" are likely to achieve this goal.

First, we make the case that there already exists a very useful prototype of an "exchange setting," which has operated for decades: Medicaid Managed Care (MMC). Over the past thirty years, most state Medicaid programs have transitioned from traditional FFS models to MMC systems.[1] Like clients in the ACA exchanges, MMC enrollees have a choice between highly regulated, capitated, competing insurance plans, but no option to join a state-run FFS program. Second, we provide evidence on how high- and lost-cost enrollees fare in these existing exchange settings, documenting large increases in health disparities by *ex-ante* health status associated with the switch from FFS to MMC. Third, we offer a new model of risk-selection in exchange settings that can rationalize our empirical results.

The existing literature on risk-selection in insurance markets is rich, but provides little evidence that directly applies to exchange settings. For example, in Medicare, private Medicare Advantage (MA) plans compete alongside a state-run FFS program, and researchers have examined whether MA plans are

[1] According to the Kaiser Family Foundation, while only about 10 percent of Medicaid recipients were enrolled in managed care in the early 1990s, 74 percent are today.

able avoid high-cost individuals by directing them to the FFS plan.[2] However, this type of risk-selection is irrelevant for exchanges, where no FFS option exists. Similarly, many papers have found that in the so-called "wild west" of the pre-ACA individual insurance market, private plans risk-select by simply denying coverage to high-cost enrollees, charging them higher premiums, or carving out coverage of pre-existing conditions.[3] But the ACA's guaranteed issue, community rating, and minimum credible coverage mandates make these blunt risk-selection strategies obsolete. Like the ACA, MMC clients choose exclusively among competing, private, capitated plans that must provide a minimum package of benefits, accept all eligible enrollees, and abide by community rating (in the MMC case, with a universal premium price of $0). For these reasons, as well as others documented in Section 2, we believe MMC offers researchers and policy-makers the best model for understanding how risk-selection in the ACA insurance exchanges is likely to evolve.

Our empirical work requires comparing the outcomes of *ex-ante* high- and low-cost individuals in settings where the state is and is not the residual claimant on enrollees' health costs. As such, we need an exogenous *ex-ante* marker of costs that is unrelated to the quality of care an individual might receive (e.g., being born low-birth-weight would not necessarily qualify as exogenous). We also require a policy change whereby the risk of covering an enrollee's health costs shifts from the state to a private plan.

We argue that Texas' county-by-county transition from FFS to MMC is a suitable natural experiment, and focus on changes in the infant health outcomes of children of U.S.-born black and Hispanic mothers (and use foreign-born black and Hispanic mothers, who generally did not qualify for Medicaid during our sample period, as placebo groups). In Texas, these children make up the majority of Medicaid births, but have very 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 His-

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

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

panic mothers.[4] 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 Hispanics, yet MMC plans receive the same capitation payments for the two groups.

Having identified our *ex-ante* high- and low-cost groups, we use detailed birth records data to explore how their outcomes change after the switch from FFS to MMC. Mortality rates for children 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 rate 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 children of U.S.-born mothers are driven by changes in policy, and not by coincident changes in unobserved, demographic-specific factors related to health.

With respect to quality of care, we 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 a private physician's office). These results suggest that black women experience a substantial decline in access to care and providers during their pregnancies, while Hispanic women do not.

Finally, given that insurers find it unprofitable to cover them and provide them with worse care, we examine whether high-cost women respond by changing their fertility after MMC.[5] We document that the county-by-county switch to MMC leads to a significant decrease in births to U.S.-born black women, concentrated among unmarried women. Data limitations prevent us from apportioning this decrease among miscarriages, abortion, or fewer conceptions. In summary, under MMC, infants whose costs are likely to exceed the capitation payment die more frequently, experience worse health outcomes, or

[4] This black-Hispanic gap in health has also been widely documented in other settings. We review this evidence in Section 6.

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

are not born at all.

To rationalize our empirical findings, the final part of the paper presents a simple dynamic model on risk-selection in exhange settings. Each period, profit-maximizing plans choose the level of care to provide to high- and low-cost patients (i.e., those whose expected costs are above and below the capitation payment, respectively). 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 themselves self-select (though they may engage in such tactics as well). Instead, they can learn about patient costs and profitability over time and adjust the quality of care accordingly based on whether they wish to retain the patient.

We can only offer qualitative evidence (from plan websites) that plans tailor benefits in this manner, as Texas Medicaid does not systematically collect any data on benefits. As such, we do not argue that our model is the only way to rationalize the results we have found, though we hope it can serve as a useful starting point to future researchers studying risk-selection in insurance exchanges. It is important to emphasize, however, that the empirical evidence does argue against perhaps the most obvious alternative model—that capitation over-incentivizes cost-control relative to FFS and thus leads to a universal deterioration of care, which differentially hurts high-risk types. Such a model would predict either small negative or no effects on (low-risk) 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.

While we believe its similarity to the ACA insurance exchanges is an important and overlooked aspect of MMC, better understanding how MMC operates is critical in its own right. By 2019, it will serve as the primary insurer for roughly 32 million individuals (ten percent of the U.S. population).[6] Two important papers have studied the county-by-county switch from FFS to MMC in California, using 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.[7] Aizer *et al.* (2007) find that pre-natal care and birth outcomes deteriorate under MMC in California.[8] Neither paper finds evidence of risk-selection in California. As we discuss later, this discrepancy between Texas and California is consistent with differences in MMC program details between the two states. Thus, comparing our results with past work can help illustrate the trade-offs involved in MMC (and, by extension, insurance exchange) program design.[9]

[6]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.

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

[8]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.

[9]In related work, Currie and Fahr (2005) use data from the National Health Interviews Survey (NHIS), and examine how state-level MMC penetration is related to individual-level Medicaid coverage and utilization of care among *children*. They find that higher MMC penetration is associated with lower Medicaid coverage and care utilization among black children with family incomes just above the poverty line (but not among those living below the poverty line). They find no statistically significant changes in coverage or utilization

Our work also contributes to the recent literature on substitution between safety-net programs (see, e.g., Borghans *et al.*, 2012). We find 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 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.

Finally, our paper relates to the large literature on infant health disparities. A mounting body of evidence has traced the origins of adult well-being to fetal and early childhood health (see Almond and Currie, 2011 for an overview), highlighting how early-life health disparities may perpetuate economic inequality in adulthood (Currie, 2011). Additionally, several papers have documented how public safety net programs (including Medicaid) can reduce these disparities through improving the health of the most disadvantaged children (Hoynes *et al.*, 2012; Miller and Wherry, 2014; Aizer and Currie, 2014). Our results highlight the possibility that program *designs* that ignore insurer incentives may exacerbate the very disparities the program aimed to close.

The remainder of the paper is organized as follows. Section 2 describes the transition from FFS to MMC in Texas. Section 3 introduces the main data source and empirical strategy, and Sections 4 and 5 present the results. Section 6 lays out our theoretical framework. Section 7 concludes.

2 Background on Medicaid and the transition to MMC in Texas

In 1995, the Texas legislature voted to begin a staggered, state-wide shift from traditional Medicaid FFS to Medicaid managed care. The Texas Health and

among Hispanic children. Given that black-Hispanic cost differences among children are far smaller than cost differences among infants, we might not expect to find large risk-selection incentives in this setting. Unlike our paper, which focuses on the supply-side incentives of MMC plans to provide different levels of care to high- and low-cost groups, Currie and Fahr (2005) generally attribute the decline in Medicaid utilization among blacks to demand-side factors.

Human Services Commission (HHSC) set the order in which counties would switch (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, while being small enough to limit the costs related to any unforeseen transition issues. Larger urban counties switched next, and rural counties switched most recently in 2012. 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. We use this county-by-county rollout of MMC as our source of identification, and it is reassuring that the schedule was set by a central office and not negotiated by individual counties. Indeed, Appendix Table 2 shows that the rollout of MMC is uncorrelated with changes in county-level economic measures.

In Texas, as in almost all states, pregnant women and infants—the population we study—are eligible for Medicaid if their family incomes fall under 185 percent of the federal poverty line (FPL). Once managed care was implemented in a county, participation among Medicaid enrollees was mandatory. Enrollees always have at least three insurers in their county from which to choose (very similar to what Dafny *et al.* (2014) find for the ACA exchanges). The large majority (83 percent) of pregnant women make an active choice among MMC plans, suggesting an important role for plan reputation.[10] As with the ACA exchanges, undocumented immigrants are not eligible for Texas Medicaid during our sample period (even though these pregnant mothers' future U.S.-born children would be), and in fact many legal immigrants are also ineligible in

[10] The remaining 17 percent are default-enrolled into a randomly assigned plan. 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. Note that the 83 percent figure is the current level of active enrollment (we do not have default rates during our sample period). We are grateful to Stephanie Goodman at Texas HHSC for this information.

the state.[11]

MMC 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 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 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."[12]

A very 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].[13]

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. We discuss

[11]For example, as a result of federal welfare reform in 1996, most legal immigrants were subject to a five-year waiting period for Medicaid coverage during our sample period. While some states chose to extend Medicaid coverage to legal immigrants during the five- year waiting period, Texas did not. In addition, Texas denies federal Medicaid coverage to many legal immigrants even after the five-year period. See: `http://www.nilc.org/document.html?id=159`.

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

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

these discretionary services further in Section 6.2.

Finally, as we argue in Section 6, the mechanism by which plans may engage in risk-selection is by encouraging high-cost patients to switch to competitor plans. While they may do so by simply providing some patients with worse healthcare or fewer discretionary services, it is also relevant to point out that plans can actually drop patients who do not comply with their regulations. MMC plan handbooks state that clients can be dropped for reasons including not following the doctor's advice, repeated emergency room visits, and missing appointments.[14]

Before proceeding to the empirical work, it is worth discussing important differences between Texas MMC and the ACA exchanges. First, the ACA exchanges provide clients greater choice than does MMC —insurers typically provide additional coverage in exchange for higher premiums. While we suspect that this additional choice will exacerbate selection as in Cutler and Reber (1998), we cannot directly examine the effect of variation in premiums prices (which in MMC are all $0).

Second, capitation payments in the ACA exchanges will eventually be risk-adjusted prospectively: they will be based on an individual's health conditions documented in the previous twelve months, with weights calibrated using commercial data from *Marketscan*. Risk-adjustment will likely mitigate selection, though not eliminate it. In particular, it is unclear whether the *Marketscan* data reflect the costs in the exchange population, and thus the predictive power of the weights may be small. Additionally, it is difficult to gather a full year of prospective health data with an unstable enrollee population—indeed, recent estimates suggest that over 40 percent of exchange enrollees will have their coverage status switch either to Medicaid or to employer insurance in the course of twelve months (Sommers *et al.*, 2014). MMC also faces high churn (e.g., many women are only eligible when pregnant) and in fact most states do not attempt to risk-adjust capitation payments.[15]

[14]See, for example, pages 6-7 of the Parkland Community plan handbook here: `http://parklandhmo.com/Handbooks/parkland%20english.pdf`. We were unable to ascertain how often plans drop clients.

[15]See Winkelman and Damler (2007). They find that only 13 states have implemented

3 Data and empirical strategy

As noted in Section 1, we will examine how outcomes for high- and low-cost groups evolve after a county switches from FFS to MMC. We first describe our main data source, and then explain how we use the county-level rollout of MMC to identify MMC's effects on four different subgroups—low- and high-cost "treatment" groups that are both largely eligible for Medicaid, as well as low- and high-cost "placebo" groups that are largely ineligible. We then provide evidence that U.S.-born black and Hispanic pregnant women have large expected cost differences, while both having very high Medicaid coverage rates, whereas their foreign-born counterparts have similar cost differences but are generally ineligible for Medicaid.

3.1 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.

Counties switched from FFS to MMC between 1993 and 2006 (Appendix Table 1). We drop the four pilot counties that switched in 1993 as we could not determine when the pilot period ended. We also drop counties that switched into MMC in January 2006 because this time period is concurrent with the influx of black refugees following Hurricane Katrina in September 2005.[16] Therefore, we limit our sample of analysis to conceptions by mothers residing in

MMC risk-adjustment or are considering it. Also see Weiner *et al.* (2012) for a review of issues related to risk-adjustment in the ACA insurance exchanges.

[16]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.

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 observations missing information on gestation, parity, mother's age, mother's race/ethnicity, and mother's marital status, which leaves us with 2,814,681 observations.

3.2 Empirical design

Our empirical strategy is straight-forward: 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.[17] Our estimating equation thus takes the form:

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

for births by mothers residing 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 each subgroup of interest, and then test whether β coefficients vary significantly across groups. Moreover, as noted earlier, by examining not only how the *gaps* between subgroups change but also how each subgroup fares in an *absolute* sense under MMC versus under FFS, we can better sort through potential mechanisms.

[17] This method is equivalent to estimating the corresponding individual-level regression with no individual-level controls.

3.3 Selecting high- and low-cost treatment and placebo groups

In Appendix Table 3, we present summary statistics for the entire sample, as well as several demographic subsets of mothers: U.S.-born blacks, U.S.-born Hispanics, foreign-born blacks, foreign-born 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. 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 black-Hispanic infant health measures are substantial: children of U.S.-born black mothers have rates of low-birth-weight, pre-term delivery, and death that are, respectively, 71 percent, 41 percent, and 74 percent greater than the corresponding rates for children of U.S.-born Hispanics. The black-Hispanic gap is very similar among the foreign-born as well—slightly larger for mortality and low-birth-weight, and slighty smaller for pre-term births. These large differences, while perhaps striking, are completely consistent with a long public health literature. Hispanic infants in the U.S. are remarkably healthy—in fact, researchers use the term "Hispanic paradox" to describe the fact that despite socio-economic deprivation comparable to blacks, they have much better health outcomes.[18] We generally take the cost differences between blacks and Hispanics as given, though briefly review potential explanations in the footnote below.[19]

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

[19] 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., as these advantage appear to dissipate slightly in the second generation with assimilation (see Guendelman and Abrams, 1995 as well as our Table 3). 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.

Of course, what matters to health plans is expected cost above the capitation payment. In Appendix Table 4, we use Texas hospital discharge data over 1999-2004 to estimate differences in delivery and newborn costs by race and ethnicity (mother's place of birth is not included in these data), conditional on year and county effects (as capitation payments are adjusted in this manner).[20] As column (1) shows, black newborns incur charges 81 percent greater than their Hispanic counterparts, or, in absolute terms, an additional \$4,218.[21] This absolute difference in initial hospital charges substantially understates the overall cost difference between black and Hispanic infants, as the elevated medical costs of at-risk births persist well beyond the first 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).[23]

Appendix Figures 1(a) and 1(b) show the 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.[24]

---

[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 4, the cost differences from the discharge data that we report in Table 4 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 4 give a good approximation of proportional cost differences.

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

[23] Note 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.

[24] 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, the cost differences associated with black versus Mexican-origin mothers was actually slightly larger when restricted to U.S.-born members of those groups.

On the whole, Appendix Table 4 and Appendix Figures 1(a) and 1(b) suggest that blacks and Hispanics serve as good proxy groups for high- and low-cost patients. Appendix Table 4 shows that differences by mother's age are also significant, however—black and Hispanic mothers age 35 and older have delivery costs 15 percent and 14 percent greater than their younger counterparts, respectively (cols. 4 and 5). As such, births to older mothers represent a small but expensive subset of Medicaid births. Yet given that older mothers are unlikely to have a future birth, plans may be less worried about these unprofitable clients returning in the future. For this reason, our empirical work generally focuses on differences by race and ethnicity.

While we have identified *ex-ante* high- and low-cost groups, it remains to be shown that we can separate them into treatment (i.e., Medicaid eligible) and placebo (i.e., Medicaid-ineligible) groups. Texas only began collecting Medicaid status on the birth certificate starting in 2005, after our sample period. Moreover, the Medicaid variable is problematic in the context of studying MMC because privatizing Medicaid seems to have had the effect of making enrollees or providers incorrectly record some Medicaid births as being covered by a private or "other/unknown" insurer, a possibility hypothesized by Aizer *et al.* (2007). We estimate that 30 percent of Medicaid births are incorrectly recorded.[25]

Appendix Table 5 shows Medicaid coverage in 2005 for our different subsets of Texas births (means are grossed up by 1.3 to adjust for under-reporting). Medicaid covered approximately 84 and 88 percent of births to U.S.-born black and Hispanic mothers in 2005, respectively; these births accounted for 56 percent of total Medicaid births.

As noted earlier, all undocumented immigrant women (and many documented immigrants as well) are excluded from Texas Medicaid during our sample period. In the Appendix, we estimate an upper bound of 44 percent for the documented share of foreign-born Hispanic women in Texas, which is itself an upper bound on the *Medicaid-eligible* share. As such, foreign-born status should serve as an excellent marker of Medicaid ineligibility and thus

[25] See the notes to Appendix Table 5 for this calculation.

placebo group status, and, indeed, in Table 5 the Medicaid share for foreign-born Hispanics is less than a third of that for their U.S.-born counterparts. Medicaid coverage of foreign-born blacks is only forty percent of that for U.S-born blacks, but given the small sample size of immigrant blacks, immigrant Hispanics represent our more meaningful falsification group. As an additional check, we also show results for married non-Hispanic whites, who have very low Medicaid rates and thus serve as another placebo group.

Before moving on to the results, we wish to emphasize that none of the analysis that follows proves that plans specifically discriminate against African-Americans *per se*. Rather, we, as researchers, need to use a proxy (race/ethnicity) 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 be profitable for them to target race specifically.

4 Results on birth outcomes

4.1 Main results

Table 1 compares changes in mortality for children of U.S.-born black and Hispanic mothers after MMC (for ease of exposition, unless otherwise noted, "black" and "Hispanic" will refer to U.S-born black and U.S.-born Hispanic mothers, respectively). For this and many other tables in this section, each pair of columns presents first the estimate for blacks 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.[26]

[26]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.

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 births to black mothers, while falling by 0.154 percentage points or $^{0.154}/_{0.715} = 21.6$ percent among births to Hispanic mothers.[27] 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 ($p \approx 001$). 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 in Appendix Table 2 that these trends were uncorrelated with MMC rollout. 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 nearly doubles ($^{0.429}/_{(1.26-0.822)} = 97.9$ 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 1. 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 2 shows results for other birth outcomes. Again, health significantly worsens for black infants (cols. 1, 3, 5): incidence of pre-term birth (defined 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 include the sex ratio as an outcome (col. 7), given the growing literature documenting its positive correlation with maternal well-being during pregnancy (as male

[27]More precisely, our mortality measure is an indicator for whether a death certificate is matched with the birth certificate by the time we obtained our data in 2010. Thus, for births in our sample, this measure captures both infant mortality and child mortality through ages beyond the first year of life.

fetuses are more likely to miscarry).[28] The male share of births falls for black mothers, but not significantly.

The even-numbered columns showing the Hispanic results tell a very different story. While results for birth weight 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 in the table, the black-Hispanic gaps move in the direction of increasing health disparities after MMC, and are significant at the five percent level.

Appendix Figures 2 and 3 show graphically the results for pre-term and male share of births, which showed the largest black-Hispanic post-MMC divergences in Table 2. 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 2).

4.2 Robustness checks

Results for placebo groups. Appendix Tables 6 and 7 show that the deterioration of outcomes for children U.S.-born black mothers and the improvement of outcomes for children of U.S.-born Hispanic mothers do not extend to children of foreign-born mothers. The only two (marginally) significant outcomes for children of foreign-born blacks are in the opposite direction of our main U.S.-born results. For children of foreign-born Hispanics, the results are all small, statistically insignificant, and often of the opposite sign of the main U.S.-born results. The large number of foreign-born Hispanic mothers ($N > 600,000$) makes this falsification exercise particularly demanding and strongly suggests the improved outcomes of children born to U.S-born Hispanic mothers are not driven by demographic-specific health factors that change in a manner coincident with MMC implementation. Finally, Appendix Table 8 also shows no effect of the MMC transition for married non-Hispanic whites, who, as we showed in Appendix Table 5, are also unlikely to be on

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

Medicaid.

Changes in selection. Our results are consistent with blacks receiving lower-quality care relative to Hispanics after MMC, but also with negative changes in selection into birth for black versus Hispanic infants. Appendix Table 9 tests whether the incidence of maternal risk-factors changes for U.S.-born blacks versus Hispanics after MMC. Cols. (1) and (2) show that after MMC, mothers in both groups are younger on average (a result we revisit later); cols. (3) and (4) show both groups are 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 are more likely to smoke (though neither result is significant on its own). Of the three outcomes, only one (smoking) shows 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 1 and 2 is, if anything, understated.

Indeed, when we re-run regressions in Appendix Tables 10 and 11 for each of the birth outcome variables using individual-level data and controlling for all plausible pre-determined covariates on the birth certificate (see the table notes), the results are essentially unchanged. Given how stable the coefficients are with and without controls and the results on selection in Appendix Table 9, we are confident that our birth outcome results are not driven by selection.[29]

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 health care. Aizer *et al.* (2007)'s estimate of the deleterious effects of MMC in California on neonatal death (a fifty percent increase) is larger than the mortality increases we find for infants born to black mothers (11.6 percent). Like us, Aizer *et al.* (2007) find larger effects on mortality than on pre-term or low-birth-weight rates.

[29] It is also reassuring that Aizer *et al.* (2007) find that regressions with and without mother fixed effects yield similar results, suggesting little effect of selection in their California MMC setting. Texas no longer provides researchers data with mother identifiers, so we cannot compare siblings born before and after MMC.

Additionally, recent work suggests that maternal stress during pregnancy has important effects on birth outcomes, including pre-term birth and the sex ratio. For example, Lauderdale (2006) finds that women with "Arabic-sounding" names exhibited a fifty percent increase in pre-term births after September 11, 2001; Persson and Rossin-Slater (2014) find that *in utero* exposure to maternal stress due to the death of a family member leads to a 15 percent increase in the likelihood of a pre-term birth. Fukuda *et al.*, 1998 and Catalano *et al.*, 2005 find that the Kobe earthquake in Japan and the unemployment rate, respectively, increases in the male neonate death rate, resulting in changes to the sex ratio which are slightly larger than the effects we find in Table 2. While it is of course impossible to objectively compare the stress associated with (our hypothesized) decrease in care over the course of a pregnancy with the events (many one-time, acute episodes) examined in these studies, the impact we find on outcomes does not appear grossly implausible.

4.3 Results on birth inputs

The birth certificate data also 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 3 shows results for indicators of pre-natal care. The first two columns show that 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 results for other thresholds of pre-natal care initiation are not significant). 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.[30] 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.*,

[30] 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.

2012).[31]

The most striking change in pre-natal inputs we find is a shift in the share of black mothers seeking pre-natal care in a public clinic, as opposed to a hospital or a 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 4 shows, the effect is very large. Table 4 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 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, providing them more limited care, or plans contracting with fewer providers in black neighborhoods. Alternatively, MMC plans may simply be directing *all* clients to free clinics as a way to cut their costs; while the coefficient 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 or 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.

On the whole, it appears that the relative deterioration of care (as measured by the pre-natal inputs recorded on the birth certificate) for blacks relative to Hispanics is likely too small to fully explain the large increases in outcome disparities documented in Tables 1 and 2. Moreover, while we saw large absolute improvements for Hispanic outcomes (as opposed to merely improvements relative to blacks' decline) after MMC, we see little absolute improvements for Hispanic inputs. Given that we find no evidence that selection changed, this pattern of results suggests a role for unobserved aspects of care. For example,

[31] Recommended weight gain depends on pre-pregnancy BMI. According to the CDC, 57 and 64 percent of Hispanic and black women are overweight, respectively. The Institute for Medicine recommends weight gain during pregnancy of 15-25 pounds when women begin pregnancy overweight. As such, we choose cut-offs of 15 and 20 pounds in our regression analysis, as higher cut-offs would unlikely affect fetal health for this (overweight) population.

a plan may be more likely to provide a low-cost client with the in-home visits and free transportation that we noted in Section 2, to approve a visit with a favorite provider, or to arrange for the birth to take place in a hospital with a NICU (a dimension of care that Aizer *et al.* (2007) found to be important in California, but that we cannot examine as hospitals are not identified on the Texas birth certificate). We provide some qualitative evidence linking plan benefits to the demographic composition of the area it serves in Section 6.2.

5 Results on fertility

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

5.1 Main results

Table 5 presents regression results based on equation (1) with the share of births to U.S.-born black mothers in a conception county-month as 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 5 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).[32]

[32] 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.

Figure 2 (a) shows results graphically. A decrease in level and break from trend coincident with MMC's introduction can be seen. Figure 2 (b) depicts the evolution of the black unmarried share of all births, and indeed the decrease post-MMC is even more striking.

As a robustness check, we drop each county individually to ensure that no single county was driving our results—in all cases, the coefficient in col. (2) remains negative and statistically significant. We also examine whether the "echo" of this birth composition result could be seen in 2005-2011 American Community Surveys (county is not recorded in earlier years of the ACS) as the children born during our sample grew older. Appendix Figure 5 shows some evidence of a decrease in the black share of cohorts conceived after MMC, though the ACS sample size is obviously far smaller than our universe of births.

Finally, we note that the decrease in black births is not a result of changing migration patterns for blacks in Texas. 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 12, we find no such effect.

5.2 Exploring mechanisms for the fall in black births

Miscarriages. As mentioned 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, Fukuda *et al.* (1998) find that 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. If one assumes that the decrease in fertility after the earthquake came *only* through miscarriages (as proxied by the sex ratio) and not also decreases in conceptions, it would suggest our entire black fertility effect could be explained by miscarriages.[33] 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 also considered whether an increase in abortions among blacks may account for the drop in fertility. Black women may be more likely to terminate a pregnancy either in reaction to the deterioration in prenatal care under MMC or because physicians under MMC are more likely to discuss the option. In Texas, Medicaid cannot directly pay for clients' abortions, but insurers may choose to contract with physicians who are more willing to suggest it. Additionally, recall from Table 4 that blacks increasingly turn to public clinics for pre-natal care after MMC, where physicians may be more likely to discuss abortion. As Texas releases individual-level abortion data from 1998 onward, these data only enable us to study the last group of counties switching to MMC in 1999, and provide a short, at most one-year, pre-period.[34] While we attempted to estimate the impacts of MMC on abortion rates, the results were sensitive to modeling choices, which is not surprising given the short pre-period.[35]

5.3 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. However, it is worth noting that Kearney

[33]Table 2, blacks' sex ratios fell 0.627 percentage-points post-MMC. Taking the implied elasticity from the Kobe event-study, we would expect a $0.627 * \frac{6}{1.5} = 2.51$ percent decrease in fertility, which is larger than the 1.1 percent drop we found in col. (1) of Table 5. Recall, however, that our estimate for the black change in the sex ratio is imprecisely estimated.

[34]We say 'at most' because we find that the number of abortions reported climbs steeply in the first quarter of 1998, suggesting that reporting does not become complete until April.

[35]Results not reported but available upon request. In particular, results were sensitive to how we chose to define treated mothers (because most abortions take place 2-3 months after conception, one might wish to count women who conceive three months before MMC as "treated.") 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.

and Levine (2009) find, nationally, that when low-income women are provided greater access to contraception, births fall differentially for Hispanic women. This evidence suggests that they, not blacks, have greater unmet demand for contraception. Moreover, an increase in black abortions would 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 higher-cost Medicaid subgroups such as older mothers differentially had the greatest unmet demand for contraception under FFS. Recall that in Appendix Table 9, we saw that the share of births to mothers over 35 fell 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 as well as married whites (see col. 7 of, respectively, Appendix Tables 6, 7 and 8).

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. Comparing our results with past work on California MMC nicely illustrates the classic trade-off between incentivizing cost-control versus preventing risk-selection. In particular, California has lower-powered incentives, as expensive cases are "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 in California (Duggan, 2004). However, neither paper finds evidence that high-cost births went "missing" after MMC or that health disparities widened, as we do in Texas. Thus, carving out high-cost cases appears to increase total spending, but also guards against incentives to avoid or discourage high-cost births.

6 Modeling risk-selection in "exchange" settings

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. Below we present a simple model of insurer incentives in exchange settings such as MMC. As we only have data on individuals and not the activities of plans themselves, the connection between the model and the empirical results is not direct—at best, we can say that individual outcomes evolve in a manner consistent with the actions we hypothesize plans take.

6.1 Modeling incentives in exchanges 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 (defined broadly; in our context it could include factors like the number of pre-natal visits and the quality of the hospital at which the mother will deliver) and those associated with outcomes $c_i(\theta)$, where c varies by patient type.[36] 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 (as we showed earlier, empirically there is very limited variation for costs related to the mother, and almost all variation comes from costs related to the infant).

Incentives in exchanges. In an exchange setting such as 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 next period. In our MMC 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 $\lambda'' < 0$. (In fact, in Texas, Medicaid recipients can change plans in the middle of a pregnancy, though we were

[36] 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.

unable to determine how frequently such a transition occurs.) 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 *relative* to those for sick clients. As we discuss in the conclusion, welfare implications are uncertain, as there are no predictions for *absolute* levels of care (which will in general depend on the relative values of ρ and p), just for the distribution of care across types. However, the fact that we do find absolute improvements for our healthy group (Hispanics) and absolute deterioration for our sick group (blacks) is very useful in falsifying another, plausible model. Capitation is often believed to over-incentivize cost-cutting, thus reducing quality of care for everyone, with potentially larger effects for the sick. The large gains for Hispanics are consistent with our model, but not the general cost-containment scenario.

Additionally, this model implicitly predicts that the effective price of child-bearing 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.[37] Moreover, if the continuation proba-

[37] 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

bility 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.

On many dimensions, the model of course simplifies how MMC or exchanges more generally work. For example, 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 (through, e.g., recommendations to friends and family).[38]

Moreover, we assume that utilization of care, health outcomes, and client retention all 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 Aizer *et al.* (2007) find that MMC-induced changes are large enough to affect birth outcomes), or that the factors that encourage retention (e.g., timely returning of phone calls) are irrelevant to health. Finally, as noted in Section 2, a more general model would allow clients to pay extra for additional coverage. These caveats notwithstanding, we hope this basic model might serve as a useful starting point for future work.

6.2 Plausibility of the model: Additional evidence on plan services from insurer websites

As documented in Section 2, MMC plans in Texas are given discretion over to whom they offer so-called "value-added services" (and of course discretion

would be pro-natalist.

[38]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.

along the many aspects of care unobserved by the state). Individual-level access and use of such services is not recorded, so we take a second-best approach and ask whether plans that operate in areas with more Hispanic clients appear to advertise more generous services (θ in the parlance of our model).

Appendix Table 13 provides plan-level data on both the demographics of the area they serve as well as the services they advertise on their website. Many plans (e.g., Amerigroup) operate across the state, and thus their demographics reflect the state's average. Others are more local and thus provide greater variation. The black/Hispanic ratio varies from 1.2% (Driscoll Children's, which serves counties near the Mexican border) to 58.1% (Parkland Community, which serves Dallas). Examples of value-added services targeted toward pregnant mothers include free baby showers, prenatal classes, gifts, in-home visits, and free transportation.[39]

By way of example, consider the services advertised by Driscoll (which serves a mostly Hispanic area) to those advertised by Parkland (which serves a mostly black area). Driscoll offers MMC clients free eyeglasses, cell phone minutes, transportation to appointments, dental care, gift cards, and a bilingual prenatal class for pregnant mothers ("Cadena de Madres"). The prenatal class includes three baby showers, baby gifts, and access a nurse and social worker (See Appendix Figure 6).[40]

While Driscoll prominently advertises its "extras," Parkland's website does not even list its value-added services; a list of these services is only found in the member handbook.[41] Parkland does not offer any of the extra services offered by Driscoll. For pregnant women, the only value-added service is a gift for completing a prenatal education class, but Parkland does not host its own course, instead offering to subsidize outside classes.

While the aggregate nature of these plan-level data is limiting, we perform

[39]Value added services are paid for by the plan, not the Medicaid program.

[40]Source: `http://www.dchpkids.com/star/services.php` and `http://www.dchpkids.com/services/?location=cadena_de_madres`.

[41]Source: `http://parklandhmo.com/healthfirst%20page.html` and `http://parklandhmo.com/Handbooks/parkland%20english.pdf`, value-added services are listed on page 9.

the basic exercise of comparing the black/Hispanic ratio in areas covered by generous plans relative to the average black/Hispanic ratio in the state (28.1 percent). Appendix Table 14 shows that the average black-Hispanic ratio is lower among plans offering baby showers (25.6%), prenatal or postnatal gifts (22.8%, 22.6%) or prenatal classes (25.1%). That is, plans serving relatively more Hispanic clients appear to offer more discretionary services. As these services are provided free of charge to the state, it appears that some of the surplus plans gain from enrollees whose costs are well below the average expected cost gets passed back to Hispanic mothers in this fashion.

Overall, our qualitative examination of plan websites and handbooks lends plausibility to the hypothesis that these plans can selectively target care across high- and low-cost patients, and that patients might respond to their care by switching plans. It is worth noting that in addition to *where* services are offered, the *types* of services offered can be used to select low-cost Medicaid eligibles; for example, prenatal classes in Spanish would only appeal to Hispanic clients, and carseats, a frequently offered postnatal gift, would only appeal to the higher-income clients who have cars.

We certainly do not view these patterns as definitive proof of plan selection. Moreover, we suspect that the most effective selection mechanisms operate at the individual level along unobserved dimensions. The inability to more directly observe mechanisms 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.[42] Hopefully, future work can make further progress in this area.

7 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 cap-

[42]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.

itated private insurance plans but does not itself administer 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 61, 13, and 41 percent, respectively, after a county switches from FFS to MMC. Quality of pre-natal care generally improves for Hispanics relative to blacks, and black birth rates fall substantially after MMC.

We explain our empirical findings through a simple dynamic model of risk-selection in "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 relative to the latter.

While we believe that our results provide compelling evidence for how care and outcomes diverge for high- and low-cost groups under MMC, their welfare implications are complicated. Given the larger number of Hispanics than blacks in Texas, average birth outcomes do not decline. However, if society wishes to shrink health disparities, then MMC may be inferior to FFS as it transfers health resources *away* from the sick to the healthy. 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.

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 unmet demand or if they are actively discouraging births among mothers because of the expected costs.

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 externality problem would not exist with a monopolistic insurer (though other problems associated with monopoly would likely arise). Our results suggests that competition may undermine the underlying policy goal of capitation in insurance—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 competitors.

With Medicaid Managed Care, the ACA exchanges, Medicare Part D, and recent calls to transform Medicare into a private premium-support program, the U.S. is moving rapidly toward providing public health insurance through a model of competing, capitated private insurance plans. While our paper points out a new challenge associated with this model, past work has warned of potential increases in costs that come with insurers losing monopsony bargaining power over providers and consumers' cognitive overload from choosing among a large set of options.[43] 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.[44] Given the direction of U.S. health policy, future work to better assess these trade-offs is of growing importance.

References

Abaluck, J. and Gruber, J. (2011). Choice Inconsistencies among the Elderly: Evidence from Plan Choice in the Medicare Part D Program. *The American Economic Review*, **101** (4), 1180–1210.

Aizer, A. and Currie, J. (2014). The intergenerational transmission of inequality: Maternal disadvantage and health at birth. *Science*, **344** (6186), 856–861.

—, — and Moretti, E. (2007). Does Managed Care Hurt Health? Evidence from Medicaid Mothers. *The Review of Economics and Statistics*, **89** (3), 385–399.

Albanesi, S. and Olivetti, C. (2010). *Maternal Health and the Baby Boom.* Working Paper 16146, National Bureau of Economic Research.

Alexander, G. R., Kogan, M., Bader, D., Carlo, W., Allen, M. and Mor, J. (2003). US Birth Weight/Gestational Age-Specific Neonatal Mortality: 19951997 Rates for Whites, Hispanics, and Blacks. *Pediatrics*, **111** (1), e61–e66.

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

[44]See Dafny *et al.* (2013) and Dafny *et al.* (2014).

ALMOND, D. and CURRIE, J. (2011). Human capital development before age five. In O. Ashenfleter and D. Card (eds.), *Handbook of Labor Economics*, vol. 4, Elsevier, pp. 1315–1486.

BAICKER, K. and DOW, W. H. (2009). Risk Selection and Risk Adjustment: Improving Insurance in the Individual and Small Group Markets. *Inquiry*, **46** (2).

BATATA, A. (2004). The Effect of HMOs on Fee-For-Service Health Care Expenditures: Evidence from Medicare Revisited. *Journal of Health Economics*, **23** (5), 951 – 963.

BAUHOFF, S. (2012). Do health plans risk-select? An audit study on Germany's Social Health Insurance. *Journal of Public Economics*.

BORGHANS, L., GIELEN, A. C. and LUTTMER, E. F. (2012). *Social Support Substitution and the Earnings Rebound: Evidence from a Regression Discontinuity in Disability Insurance Reform*. Tech. rep., National Bureau of Economic Research.

CATALANO, R., BRUCKNER, T., ANDERSON, E. and GOULD, J. (2005). Fetal Death Sex Ratios: A Test of the Economic Stress Hypothesis. *International journal of Epidemiology*, **34** (4), 944–948.

CURRIE, J. (2011). Inequality at birth: Some causes and consequences. *The American Economic Review*, **101** (3), 1–22.

— and FAHR, J. (2005). Medicaid managed care: effects on children's medicaid coverage and utilization. *Journal of Public Economics*, **89** (1), 85–108.

CUTLER, D. and REBER, S. (1998). Paying for health insurance: The trade-off between competition and adverse selection. *Quarterly Journal of Economics*, **113** (2), 433–466.

DAFNY, L., DUGGAN, M. and RAMANARAYANAN, S. (2012). Paying a Premium on Your Premium? Consolidation in the US Health Insurance Industry. *The American Economic Review*, **102** (2), 1161–1185.

—, GRUBER, J. and ODY, C. (2014). *More Insurers Lower Premiums: Evidence from Initial Pricing in the Health Insurance Marketplaces*. Working Paper 20140, National Bureau of Economic Research.

—, HO, K. and VARELA, M. (2013). Let Them Have Choice: Gains from Shifting Away from Employer-sponsored Health Insurance and Toward an Individual Exchange. *American Economic Journal: Economic Policy*, **5** (1), 32–58.

DOMINGUEZ, T. P. (2008). Race, Racism, and Racial Disparities in Adverse Birth Outcomes. *Clinical obstetrics and gynecology*, **51**, 360–70.

DUGGAN, M. (2004). Does contracting out increase the efficiency of government programs? evidence from medicaid hmos. *Journal of Public Economics*, **88** (12), 2549–2572.

— and HAYFORD, T. (2011). *Has the Shift to Managed Care Reduced Medicaid Expenditures? Evidence from State and Local-Level Mandates*. Working Paper 17236, National Bureau of Economic Research.

Edgman-Levitan, S. and Cleary, P. D. (1996). What information do consumers want and need? *Health affairs*, **15** (4), 42–56.

Fuchs, V. R. (2004). More Variation in Use of Care, More Flat-of-the-Curve Medicine. *Health Affairs*, **23**, VAR–104.

Fukuda, M., Fukuda, K., Shimizu, T. and Moller, H. (1998). Decline in Sex Ratio at Birth after Kobe Earthquake. *Human Reproduction*, **13** (8), 2321–2322.

Giapros, V., Drougia, A., Krallis, N., Theocharis, P. and Andronikou, S. (2012). Morbidity and mortality patterns in small-for-gestational age infants born preterm. *Journal of Maternal-Fetal and Neonatal Medicine*, **25** (2), 153–157.

Glazer, J. and McGuire, T. G. (2000). Optimal Risk Adjustment in Markets with Adverse Selection: An Application to Managed Care. *American Economic Review*, **90** (4), 1055–1071.

Grossman, M. (1972). On The Concept of Health Capital and the Demand for Health. *The Journal of Political Economy*, **80** (2), 223–255.

Guendelman, S. and Abrams, B. (1995). Dietary Intake Among Mexican-American Women: Generational Differences and a Comparison with White Non-Hispanic Women. *American Journal of Public Health*, **85** (1), 20–25.

Haywood L Brown, Y. J., Monique V Chireau and Howard, D. (2007). The "Hispanic Paradox": An Investigation of Racial Disparity in Pregnancy Outcomes at a Tertiary Care Medical Center. *American Journal of Obstetrics and Gynecology*, **197** (2), 197e1 – 197e9.

Hoynes, H., Schanzenbach, D. and Almond, D. (2012). *Long Run Impacts of Childhood Access to the Safety Net*. Tech. rep., National Bureau of Economic Research.

Isaacs, S. L. (1996). Consumer's information needs: results of a national survey. *Health Affairs*, **15** (4), 31–41.

Kaestner, R., Dubay, L. and Kenney, G. (2002). *Medicaid Managed Care and Infant Health: A National Evaluation*. Working Paper 8936, National Bureau of Economic Research.

Kearney, M. and Levine, P. (2009). Subsidized Contraception, Fertility, and Sexual Behavior. *The Review of Economics and Statistics*, **91** (1), 137–151.

Langwell, K. and Hadley, J. (1989). Evaluation of the Medicare Competition Demonstrations. *Health Care Financing Review*, **11** (2), 65–80.

Lauderdale, D. S. (2006). Birth outcomes for arabic-named women in california before and after september 11. *Demography*, **43** (1), 185–201.

Leslie, J. C., Galvin, S. L., Diehl, S. J., Bennett, T. A. and Buescher, P. A. (2003). Infant Mortality, Low Birth Weight, and Prematurity among Hispanic, White, and African American Women in North Carolina. *American Journal of Obstetrics and Gynecology*, **188** (5), 1238 – 1240.

Lopoo, L. M. and Raissian, K. M. (2012). Natalist Policies in the United States. *Journal of Policy Analysis and Management*, **31** (4), 905–946.

McCormick, M. C., Bembaum, J. C., Eisenberg, J. M., Kustra, S. L. and Finnegan, E. (1991). Costs Incurred by Parents of Very Low Birth Weight Infants after the Initial Neonatal Hospitalization. *Pediatrics*, **88** (3), 533–541.

Mello, M. M., Stearns, S. C., Norton, E. C. and Ricketts, T. C. (2003). Understanding Biased Selection in Medicare HMOs. *Health Services Research*, **38**, 961–992(32).

Miller, S. and Wherry, L. (2014). The long-term health effects of early life medicaid coverage, University of Notre Dame, unpublished manuscript.

Newhouse, J. P. (2006). Reconsidering the moral hazard-risk avoidance tradeoff. *Journal of Health Economics*, **25** (5), 1005–1014.

Park, S., Sappenfield, W. M., Bish, C., Salihu, H., Goodman, D. and Bensyl, D. M. (2011). Assessment of the institute of medicine recommendations for weight gain during pregnancy: Florida, 2004–2007. *Maternal and child health journal*, **15** (3), 289–301.

Persson, P. and Rossin-Slater, M. (2014). Family ruptures and intergenerational transmission of stress, Stanford University, unpublished manuscript.

Physician Payment Review Commission (1997). Testimony of Gail R. Wilensky. Statement before the U.S. House of Rrepresentatives Subcommittee on Health. Comittee on Ways and Means, February 24, 1997.

Reichman, N. E. and Kenney, G. M. (1998). Prenatal care, birth outcomes and newborn hospitalization costs: patterns among hispanics in new jersey. *Family planning perspectives*, pp. 182–200.

— and Schwartz-Soicher, O. (2007). Accuracy of Birth Certificate Data by Risk Factors and Outcomes: Analysis of Data from New Jersey. *American Journal of Obstetrics and Gynecology*, **197** (1), 32–e1.

Rubalcava, L. N., Teruel, G. M., Thomas, D. and Goldman, N. (2008). The Healthy Migrant Effect: New Findings From the Mexican Family Life Survey. *Journal Information*, **98** (1).

Sommers, B. D., Graves, J. A., Swartz, K. and Rosenbaum, S. (2014). Medicaid and marketplace eligibility changes will occur often in all states; policy options can ease impact. *Health Affairs*, **33** (4), 700–707.

Tenovuo, A. (1988). Neonatal complications in small-for-gestational age neonates. *Journal of Perinatal Medicine-Official Journal of the WAPM*, **16** (3), 197–204.

Tommiska, V., Tuominen, R. and Fellman, V. (2003). Economic Costs of Care in Extremely Low Birthweight Infants during the First 2 Years of Life. *Pediatric Critical Care Medicine*, **4** (2), 157–163.

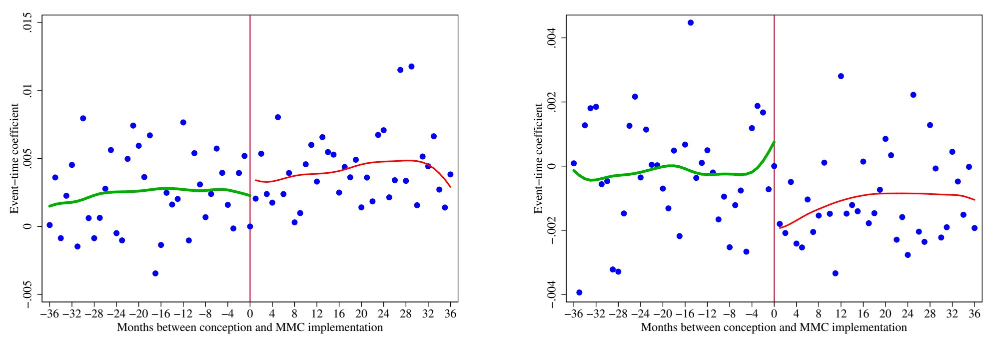
Weiner, J. P., Trish, E., Abrams, C. and Lemke, K. (2012). Adjusting For Risk Selection in State Health Insurance Exchanges will be Critically

Important and Feasible, but Not Easy. *Health Affairs*, **31** (2), 306–315.

Winkelman, R. and Damler, R. (2007). Risk adjustment in state medicaid programs. *Health Watch*, (57), 14–34.

Figure 1: Changes in mortality to children born to U.S.-born black and Hispanic mothers (note different scales)

(a) Blacks (b) Hispanics



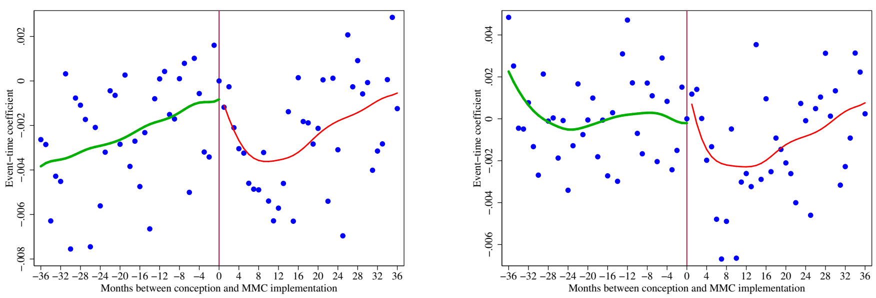
Notes: These figures show the results from estimating mortality rates for black (Figure a) and Hispanic (Figure b) births in the 36 months before and after MMC implementation. Specifically, we estimate the following equation:

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

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 2: Black share of all births before and after MMC (note different scales)

(a) Black share of births (b) Unmarried black share of births



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 1 for further details on the estimation procedure.

Table 1: 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. $^{*}p < 0.1,^{(}**)p < 0.05,^{***}p < .01$

Table 2: 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*** [0.342]	-0.607*** [0.175]	0.705* [0.376]	0.0231 [0.175]	0.906*** [0.330]	-0.146 [0.363]	-0.627 [0.412]	0.779*** [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 1 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.

Table 3: Effect of MMC on pre-natal care measures (×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.
Conceived after MMC	-2.000 [1.730]	0.0384 [0.852]	-0.0817 [0.0734]	-0.0642 [0.0697]	-2.261** [1.033]	-0.850 [0.733]	-1.276 [0.851]	0.0808 [0.689]	-2.013** [0.813]	0.163 [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 1 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 4: Effect of MMC on site of pre-natal care (×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 1 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 5: Effect of MMC on U.S.-born black birth rates

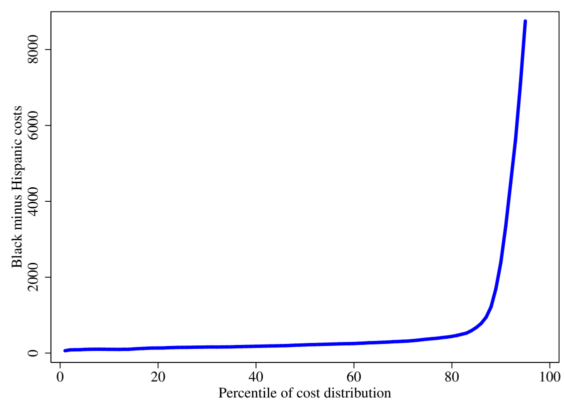
	Share of births		Log of births		
	(1) Black	(2) Bl. unm.	(3) Black	(4) Bl. unm.	(5) Non-bl.
Conceived after MMC	-0.00120	-0.00247***	-0.0309**	-0.0462***	-0.0141
	[0.000790]	[0.000630]	[0.0130]	[0.00843]	[0.00920]
Mean, dept. var.	0.105	0.0678	5.526	5.091	7.070
Reg. obs. (cells)	26021	26021	3672	3672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

Notes: See notes to Table 1 for details about the data, sample and specification. When logs are used in cols. (3) through (5), 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.

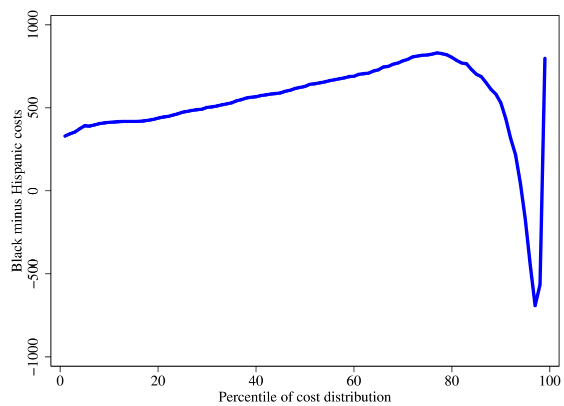
FOR ONLINE PUBLICATION

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

(a) Newborns



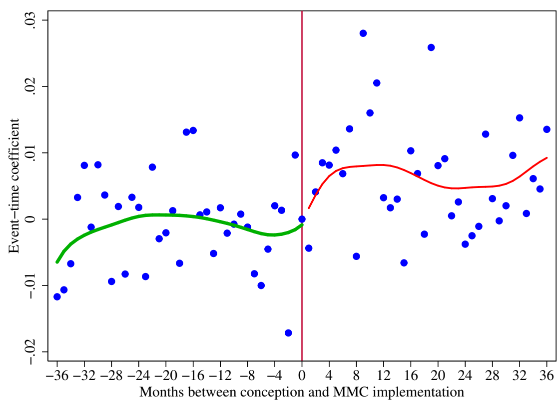
(b) Deliveries



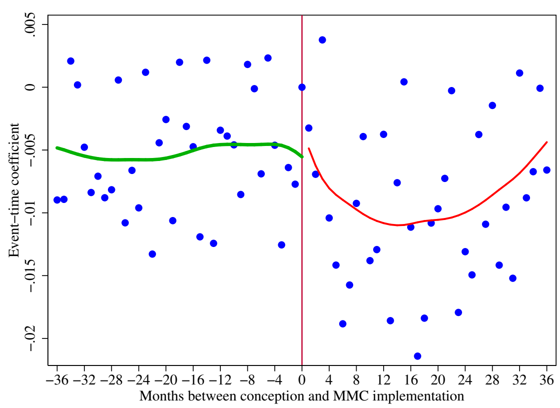
Notes: Figures are based on data from public-use Texas Hospital discharge data (see `http://www.dshs.state.tx.us/THCIC/Hospitals/Download.shtm` to download these data). For each graph, the value of the Hispanic n^{th} percentile is subtracted from the value of the Black n^{th} percentile. Because of 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.

Appendix Figure 2: Changes in pre-term share of births to U.S-born black and Hispanic mothers (note different scales)

(a) Blacks



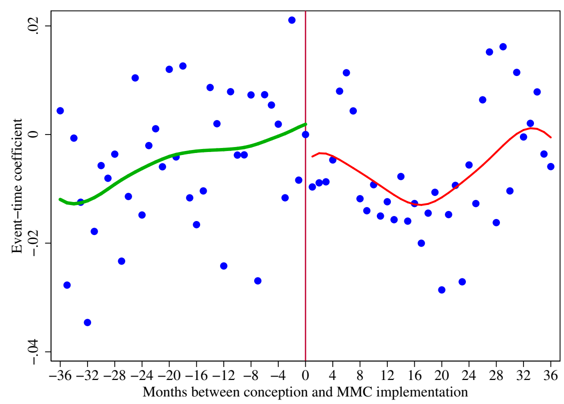
(b) Hispanics



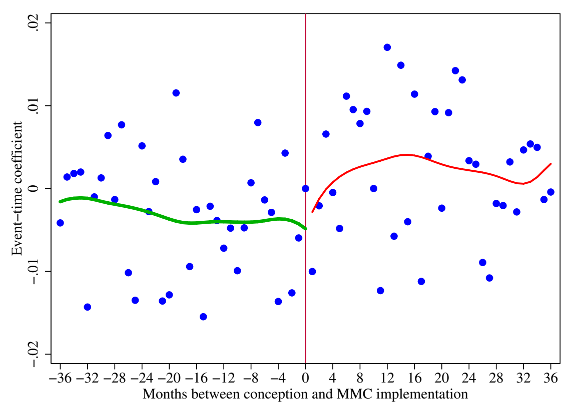
Notes: These figures are identical to those displayed in Figure 1 except that pre-term birth serves as the outcome variable.

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

(a) Blacks

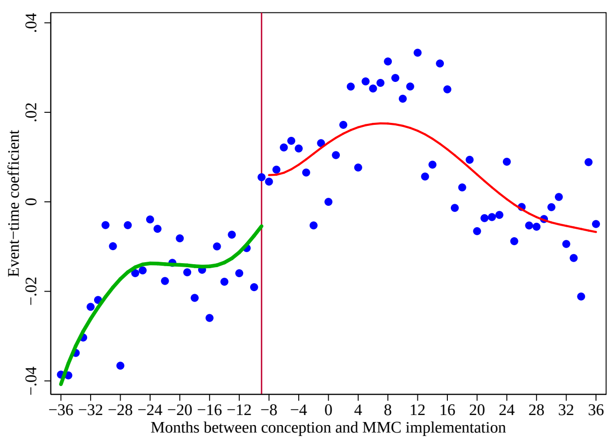


(b) Hispanics



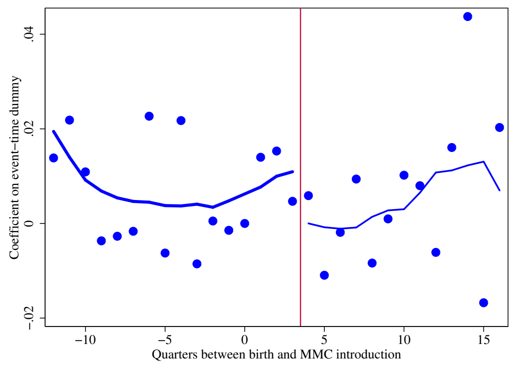
Notes: These figures are identical to those displayed in Figure 1 except that share male serves as the outcome variable.

Appendix Figure 4: Share receiving pre-natal care in a public clinic, U.S.-born black mothers



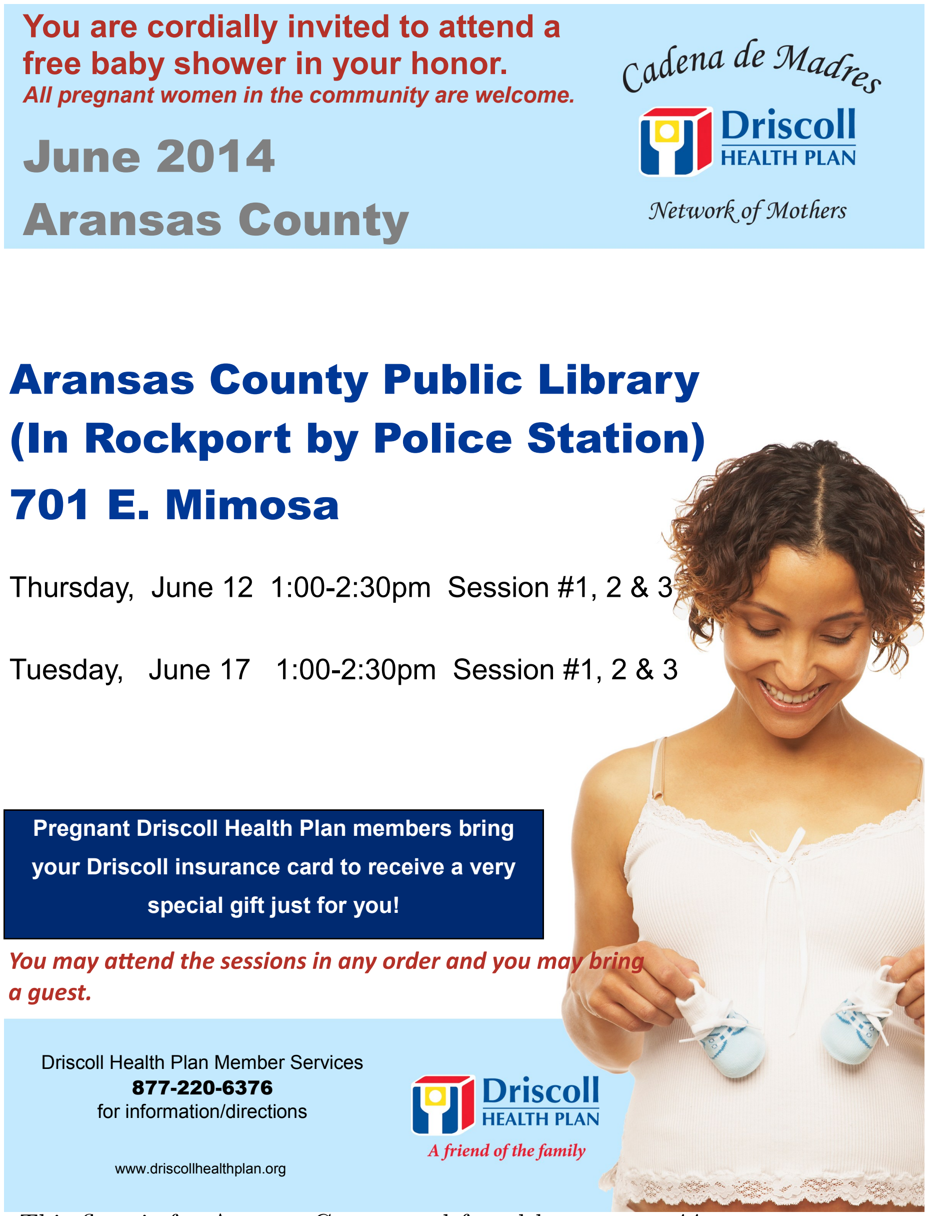
Notes: This figures are identical to those displayed in Figure 1 except that share receiving pre-natal care in a public clinic serves as the outcome variable and the pre- and post-period are defined as births (not conceptions) before and after MMC.

Appendix Figure 5: 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 Figure 6: Driscoll's "Cadena De Madres" Flyer



Notes: This flyer is for Aransas County and found here: `http://www.dchpkids.com/pdf/AransasInvite.pdf`. All flyers and other information on Cadena de Madres is found here: `http://www.dchpkids.com/services/?location=cadena_de_madres`. Note that the class is open to all pregnant women, but it is only free for those in Driscoll's MMC plan (called "Driscoll Health Plan"). Driscoll Health Plan members also receive a special gift for attending, as advertised.

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. 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.

Appendix Table 3: 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.

Appendix Table 4: 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.

Appendix Table 5: 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 3, 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. For example, comparing conceptions from 2004-2005 to those in 2007-2008 in the counties that switched to MMC in 2006, the reported Medicaid share falls from 64.7 percent to 49.9 percent (it did not fall in other counties). This drop suggests that 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. 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. We thus "gross up" the Medicaid share by 1.3 in this table

Appendix Table 6: 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	0.0305	-0.677	-0.869*	0.124	2.667*	0.522
MMC	[0.256]	[0.849]	[0.462]	[1.030]	[1.559]	[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 7: 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	-0.0424	0.234	-0.184	-0.238	-0.0727	0.000427
MMC	[0.0535]	[0.483]	[0.181]	[0.364]	[0.472]	[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 8: 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	0.0528	0.261	0.0163	-0.0874	-0.115	0.0254
MMC	[0.0443]	[0.221]	[0.156]	[0.186]	[0.198]	[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 9: Changes in risk-factors ($\times 100$) after MMC for U.S.-born black and Hispanic mothers

	Older		Diab/Hyper.		Smokes	
	(1) Bl.	(2) Hsp.	(3) Bl.	(4) Hsp.	(5) Bl.	(6) Hsp.
Conceived after MMC	-0.552** [0.238]	-0.349** [0.145]	-0.111 [0.191]	-0.247 [0.216]	-0.335 [0.254]	0.267 [0.211]
Dept. var mean	5.659	4.699	3.469	3.164	6.284	3.483
Diff/p-val	-0.203	0.398	0.136	0.593	-0.602	0.0196
Reg. obs (cells)	12832	20504	12833	20504	12808	20489
Indiv. obs.	296588	646053	296589	646053	296563	646037

Notes: See notes under Table 1 for more details about the data, sample, and specifications.

Appendix Table 10: 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 11: 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**	-0.554***	0.0415	-0.0857	0.710***
	[0.0716]	[0.168]	[0.175]	[0.399]	[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 12: 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.

Appendix Table 13: MMC Plans, Racial Composition, and Extra Services

Plan Name	Service Areas	Black Population, 2010 Census	Hispanic Population, 2010 Census	Black/Hispanic	Link to Value Added Services
Aetna	Bexar and Tarrant	480,774	1,795,499	0.2678	www.aetnabetterhealth.com/texas/members/medicaid/value-adds
Amerigroup	Bexar, Dallas, Harris, Jefferson, Lubbock, Tarrant	2,333,160	5,333,867	0.4374	www.myamerigroup.com/English/Documents/TXTX_Benefits_Overview_STAR_ENG.pdf
Blue Cross Blue Shield	Travis	133,827	555,274	0.2410	www.bcbstx.com/medicaid/star.html
Christus	Nueces	14,388	185,335	0.0776	www.christushealthplan.org/members/medicaid/value-adds
Community First	Bexar	140,382	1,136,611	0.1235	www.cfhp.com/Members/STAR/Medicaid-Mbr-Hndbk-0114.pdf
Community Health Choice	Harris, Jefferson	1,164,920	2,200,235	0.5295	www.chchealth.org/GetFile.aspx?FileId=171
Cook Children's	Tarrant	340,392	658,888	0.5166	www.cookchp.org/English/members/STAR-members/Pages/Programs-and-Benefits.aspx
Driscoll	Nueces, Hidalgo	19,987	1,633,050	0.0122	www.dchpkids.com/star/services.php
El Paso Premier	El Paso	27,419	670,604	0.0409	www.epfirst.com/PremierPlan.html
First Care	Lubbock	43,287	228,564	0.1894	www.firstcare.com/STAR_Medicaid
Molina	Dallas, El Paso, Harris, Hidalgo, Jefferson	1,842,117	5,428,123	0.3394	www.molinahealthcare.com/members/tx/en-US/mem/medicaid/star/PDF/TX_STAR_Member_Handbook.pdf
Parkland	Dallas	644,179	1,109,569	0.5806	www.parklandhmo.org/Handbooks/parkland%20english.pdf
Sendero	Travis	133,827	555,274	0.2410	www.senderohealth.com/en/members/value-adds
Seton	Travis	133,827	555,274	0.2410	www.setonhealthplan.com/members/chip/STAR_handbook-12-13%20edit.pdf
Superior Health	Bexar, El Paso, Hidalgo, Lubbock, Nueces, Travis	364,902	4,224,103	0.0864	www.superiorhealthplan.com/for-members/benefits-information/extra-benefits/
Texas Children	Harris, Jefferson	1,164,920	2,200,235	0.5295	www.texaschildrenshealthplan.org/for-members/star/
United Health	Harris, Hidalgo, Jefferson	1,170,519	3,647,950	0.3209	http://www.uhccommunityplan.com/content/dam/communityplan/plandocuments/handbook/en/TX-star-handbook.pdf

Notes: 2010 population counts by race and county are from the 2010 Census: http://www.dshs.state.tx.us/chs/popdat/st2010c.shtm. Counties served by each plan are listed here: http://www.hhsc.state.tx.us/medicaid/managed-care/mmc/STAR-map.pdf. Counties are grouped into service areas as follows: Bexar Service Area= Atascosa, Bandera, Bexar, Comal, and Guadalupe; Dallas SA = Collin, Dallas, Ellis, Hurt, Kaufman, Navarro, and Rockwall; El Paso SA = El Paso and Hudspeth; Harris SA = Austin, Brazoria, Fort Bend, Galveston, Harris, Matagorda, Montgomery, Waller and Wharton; Hidalgo SA = Cameron, Duval, Hidalgo, Jim Hogg, Maverick, McMullen, Starr, Webb, Willacy and Zapata; Jefferson SA = Chambers, Hardin, Jasper, Jefferson, Liberty, Newton, Orange, Polk, San Jacinto, Tyler, Walker; Lubbock SA = Carson, Crosby, Deaf Smith, Floyd, Garza, Hale, Hockley, Hutchinson, Lamb, Lubbock, Lynn, Potter, Randall, Swisher, Terry; Nueces SA = Aransas, Bee, Brooks, Calhoun, Goliad, Jim Wells, Karnes, Kenedy, Kleberg, Live Oak, Nueces, Refugio, San Patricio, Victoria; Tarrant SA = Denton, Hood, Johnson, Parker, Tarrant, Wise; Travis SA = Bastrop, Burnet, Caldwell, Fayette, Hays, Lee, Travis, Williamson.

Appendix Table 14: Black/Hispanic Ratio by Services Offered

Benefit	Plans offering Benefit	Average Bl/Hisp
Baby shower or other social event	Community First, Community Health Choice, Driscoll, Firstcare, Texas Children's, Superior, United	0.2559
Prenatal gifts	Aetna, BCBS, Christus, Driscoll, El Paso, FirstCare, Parkland, Texas Children's, Sendero, Seton, Superior	0.2279
Postnatal gifts	BCBS, Community First, Community Health, Driscoll, El Paso, FirstCare, Molina, Sendero, United	0.2264
Hosts prenatal classes	Amerigroup, Community First, Community Health, Driscoll, El Paso, Superior, Texas Children's	0.2513
In-home visits for high-risk members	BCBS, Christus, Seton, Texas Children's, United	0.2820

See notes to the previous table.

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.[45] According to the Pew Hispanic Center, there were 1.1 million undocumented immigrants in Texas in 2000.[46] 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.[47] 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.[48] 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.

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

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

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

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