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Is there Evidence Against the Induced Demand Hypothesis? Explaining the Large Reduction in Cesarean Rates
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Abstract:

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The induced demand model postulates that physicians respond to adverse income shocks by electing to perform more remunerative procedures. Recent work verifies the predictions of this model, finding a strong shift away from natural deliveries to the more highly reimbursed cesarean delivery in response to exogenous reductions in the fertility rate between 1970-1982. In light of this, the 10.1 percent reduction in fertility rates and contemporaneous 13.3 percent decline in cesarean rates between 1989-1996 appears to contradict the induced demand hypothesis. This paper reconciles the observed phenomenon by examining the role of the dramatic growth in managed care activity in this time period. We argue that while the inducement effect of declining fertility was pervasive in this period, its effect on cesarean rates was offset by the dramatic growth in managed care beginning in the late 1980s. Central to our argument is the removal of physician financial incentives in delivery choice under managed care. Using county-level data for a subset of this time period and instrumental variables estimators, we find a strong negative association between managed care growth and cesarean rates that persists despite controls for fertility, birth severity, risk factors, demographics, state fixed effects and hospital characteristics.

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1 Introduction

There is widespread belief and some empirical evidence that a high fraction of cesarean-section deliveries result from the demand inducing behavior of obstetricians/gynecologists (henceforth ob/gyns). Theories of induced demand suggest that physicians derive utility from income and leisure but disutility from inducement, perhaps as a result of ethical considerations or reputation effects; e.g., Dranove (1988). A straightforward implication of these models is that when income is tailored to specific medical procedures, physicians will abuse their role as informed agents to their patients by inducing demand for the more remunerative procedure until its marginal benefit equalizes the associated marginal costs.

Induced demand raises obvious ethical issues, but is also central to the health care reform debate. If there is little evidence of the benefits of alternate, more expensive medical technologies relative to their costs it bears upon health policy to evaluate and realign the financial incentives facing physicians with alternate compensation schemes. In recent years, various policies to restructure the reimbursement of deliveries and reorganize the practice environment for obstetrics have been widely discussed; see e.g., Myers and Gleicher (1993), CDC (1993). The evaluation of such policies rests on evidence that quantifies the links between both obstetrics' delivery choices and their practice environment on one hand, and the links between neo-natal/maternal health outcomes and the practice environment on the other. This paper provides some evidence in this context.

The delivery of babies has been a natural testing ground for induced demand models, most recently by Gruber and Owings (1996); see also Stafford (1990), Dranove and Wehler (1994). In the early 1990s cesareans not only reimbursed ob/gyns \$611 more on average than a natural delivery (CDC, 1993), they also required less time on average allowing an ob/gyn higher utility for inducing demand for cesareans until the marginal utility from a higher income and a reduced work load equalized the marginal disutility of inducement.[1] Over the 1965-1985 period cesarean deliveries in the United States rose from just 5 percent to 22 percent of all deliveries, while U.S fertility rates contemporaneously fell over 13 percent (National Center for Health Statistics (NCHS), 1994). Gruber and Owings (1996) conclude that it was this exogenous source of income-

[1] Although cesareans on average cost \$7826 versus \$4720 for a natural delivery in 1991, the difference in ob/gyn fees is only \$611 because of more expensive inputs in cesarean delivery. Another factor that may be of importance in deciding between a cesarean-section versus a vaginal delivery is that cesareans can be scheduled, offering the ob/gyns a non-pecuniary benefit from electing a cesarean delivery. Raw means in my data shows that amongst cesarean deliveries, an average of 35.7 percent more babies were born on Friday than on another day of the week suggesting that ob/gyns are scheduling a larger fraction of cesarean deliveries just prior to the weekend (see Table 1).

pressure that led ob/gyns to substitute towards the more highly reimbursed cesarean deliveries, driving up the cesarean-section rate.

At the outset, the decline in cesarean rates between 1989 and 1996 would appear to contradict the induced demand model. Figure 1 shows that in this period U.S. fertility rates fell 10.1 percent heralding a negative financial shock to ob/gyns, but concurrently U.S. cesarean rates *fell* 13.3 percent contradicting the central premise of the induced-demand hypothesis. This paper attempts to reconcile the observed phenomenon by examining the role of the contemporaneous growth in managed care activity. Shown in Figure 1, Health Maintenance Organizations (HMOs) grew from serving 8 percent of the health insurance industry in 1986 to 25.3 percent in 1996 (HIAA, 1996). We argue that while the inducement effect of declining fertility was pervasive in this period, its effect on cesarean delivery rates was offset by a corresponding reverse effect on cesarean rates due to the dramatic rise in managed care activity beginning in the late 1980s.

The simple premise underlying this argument is the capitation of physician fees in managed care; i.e., reimbursement fees are fixed and determined prospectively, typically independent of the volume of medical service provided. In the context of deliveries managed care reimburses cesareans and natural deliveries equally, permitting a managed care ob/gyn no direct pecuniary benefit from inducing demand for a cesarean.[2] There is also evidence that managed care directly or implicitly imposes high penalties on physicians who consistently violate the cost conservative practice patterns preferred by managed care (Robinson and Steiner, 1998). It is plausible therefore that by severing the financial link between delivery and income and altering the practice environment facing ob/gyns, the dramatic growth in managed care activity in the late 1980s and early 1990s played a role in the large decline in the national cesarean rate.

In testing for a causal role of managed care activity on the cesarean rate, we examine the empirical link between HMO market shares (measured as the fraction of a local market's population enrolled in HMO insurance) and the local cesarean rate, conditional on fertility, hospital size, risk factors, demographics and other determinants of the cesarean rate. Due to data limitations, this study will focus on the 1991-1992 time period in which cesarean rates fell 8.5 percent from its highest rate in 1989, fertility declined by 3 percent, and HMO activity contemporaneously grew over 15 percent.

[2] Of course there may be other nonpecuniary benefits such as additional leisure and scheduling conveniences that may lead managed care ob/gyns to induce demand on the margin; however, one might expect that the inducement effect arising from these factors to be second order relative to the inducement effect arising from higher income. Also, while cesareans on average take a shorter time to perform than deliveries with extended labor, it is plausible that physicians do not measure leisure purely in time input but take into account the stress of cesarean surgery and the exposure to legal risk from performing cesareans; see, e.g., Keeler and Brodie (1993). In this case managed care ob/gyns may not even receive much nonpecuniary benefit from inducing demand.

We use information about HMO activity at the county level collected by the American Association of Health Plans (AAHP); the remaining data are primarily from the National Center for Health Statistics (NCHS) which documents detailed parental, hospital and birth characteristics of each delivery in a birthing center in the U.S. These data are aggregated to the county level to match the unit of observation of the HMO data. We are aware of the possible endogenous determination of HMO market shares, and our empirical analysis addresses this issue using large across-state variations in laws that regulate the entry, expansion and operation of managed care. Our results are driven largely by the tremendous across-county variation in managed care activity, although the use of two years' data also gives us a small degree of across-time variation.

We find strong evidence that cesarean rates are negatively correlated with the growth in HMO activity. This relationship persists despite controls for risk factors, birth severity, hospital size, region, income and demographic factors, but attenuates somewhat when controlling for female labor force attachment. In particular, we find that fertility continues to exert an induced demand for cesareans, but that its effect is swamped out by the opposite effect from managed care. This result supports the hypothesis that altered financial incentives introduced by managed care played a role in the reduction of cesarean rates. Our results suggest that a 1 percentage point increase in HMO market shares is associated with a decrease of between 4 and 6 percent in cesarean rates at 1992 cesarean levels. The paper also assesses whether increased managed care activity in 1991-1992 is associated with tandem declines in quality of births. Focussing on neo-natal infant mortality, Apgar scores and birth injury/complication rates we find no such evidence.

The paper proceeds by describing, in Section 2, the trends in health insurance markets, managed care activity and cesarean births, also describing the mechanisms that lend a causal role for managed care activity in the reduction of the cesarean rate. Next in Section 3 we present a theoretical framework to examine the effects of changing HMO market shares on cesarean rates, summarize its main implications, and review the related literature. Section 4 describes the data, the empirical methodology, and the empirical results. Conclusions follow.

2 Health Insurance Markets and Cesarean Sections

2.1 Background and Trends

During the 1980s managed care enrollments grew rapidly from less than 10 million in 1980 distributed over only 30 HMOs, to over 40 million members by 1992 distributed over 230 HMOs (HIAA, 1996). The number of enrolled members represented 25 percent of the insured population in 1992, and dipped slightly thereafter (AAHP, 1994). Over this time period, a number of organizations that combined features of both traditional indemnity insurance and traditional HMOs, e.g., Point of Service (POS) plans, Preferred Provider Organizations (PPOs) and Independent Practice associations (IPAs) also emerged, in which 22.26 percent of the population was enrolled by 1994 (AAHP, 1994). Since many of the newer organizational structures blur the original distinction between managed care and traditional indemnity insurance, only those organizations listed specifically as HMOs are used in the definition of managed care in the current work, and the terms "HMO" and "managed care" are used synonymously.

What led to this dramatic growth in HMOs? In contrast to the voluminous literature on the consequences of growing managed care activity,[3] little academic attention has been given to the causes behind the escalation of managed care activity in the 1980s. It is perceived that rapid cost inflation in Fee-for-Service (FFS) health care in the 1980s renewed customers' attention in alternative systems of managed care; see Robinson and Steiner (1988), Glied (2000). Managed care permitted healthier patients to self select into contracts with fewer goods and lower prices, while simultaneously shifting risk onto providers who now faced capitated reimbursements. On the supply side, managed care units responded to new consumer demands and new market opportunities, benefiting from cream skimming the healthiest patients (Newhouse, 1984). Perhaps encouraged by the method in which managed care claimed to control costs while maintaining quality, new federal and state "enabling" laws in the 1980s further spurred the growth in managed care activity.

Concurrent with the rapid growth in HMOs in the 1980s was a corresponding decrease, following a fall in the rate of increase, in the national cesarean-section rate. Figure 1 maps the cesarean and fertility rates from 1980 to 1996. Between 1980 and 1989 cesarean deliveries rose by over 50 percent, from 16.2 percent of all deliveries to 24.7 percent of the total (NCHS, 1996). Conditional on the fact that cesareans maintained a 40 percent reimbursement premium over natural deliveries in this period (HIAA, 1996), declining fertility is thought to be one of the most compelling arguments for the rise in cesareans in this period (see Gruber and Owings, 1996); other explanations for the large increase in cesarean rates include the introduction and growth of fetal-monitoring (Shiono et. al., 1987) and the threat of malpractice suits (Localio et. al., 1993). Of interest in this study are the rates after 1989, a period over which cesarean usage gradually decreased for the first time since its widespread introduction in the early 1960s.

Of the four symptoms that account for over 80 percent of cesareans, there have been declines within diagnosis as well in this period. Traditionally a leading cause of cesareans has

[3]See, e.g., Frank and Welch (1985), McClaughlin (1988), Tussing and Wojotowycz (1994), McCloskey, Petitti and Hobel (1992), Baker and Brown (1997), Cutler, McClelland and Newhouse (1997), Feldman and Scharfstein (1999), Cutler and Reber (2000), Glied (2000).

been repeat-provider behavior, an observed phenomenon that the likelihood of a cesarean for a woman previously c-sectioned is extremely high. Repeat cesareans fell from over 9 percent of all births in 1989 to 7.5 percent in 1996, coincidental with the sudden growth in vaginal births after cesareans (VBACs) in this time period (CDC, 1998). A second symptom of cesareans is a breech or abnormal presentation, which largely precludes natural deliveries unless the fetal position is corrected prior to the second trimester. In 1989, over 90 percent of breech babies were delivered by cesarean, and this rate had subsequently fallen to 84.7 percent by 1996, although the incidence of breech had remained fairly stable (NCHS, 1996). Other symptoms favoring cesareans are decidedly subjective diagnoses: fetal distress and non-progressive labor or dystocia, which is closely related to maternal age and health. Cesareans related to both these symptoms declined continuously, albeit gradually, after 1989 (NCHS, 1996).

2.2 Mechanisms that link HMO growth with Cesarean Rates

By i) severing the link between mode of delivery and ob/gyn income, ii) selecting cost conservative providers, and iii) encouraging the use of preventive diagnostic tests (e.g., prenatal tests) to reduce ex-post costly surgery, managed care radically changed the practice environment facing ob/gyns affiliated with HMOs. In the early 1990s, HMOs also began to impose practice guidelines on participating ob/gyns, and some evidence suggests that ob/gyns whose practice patterns consistently violated such guidelines suffered high penalties including the termination of their contracts; see, e.g., Schauffler and Rodriguez (1993), Robinson and Steiner (1998). Standard theories of induced demand imply that capitated reimbursements for natural and cesarean deliveries would have first order effects on the demand inducing behavior of managed care ob/gyns, while the periodic utilization reviews would serve to reinforce the effects of this financial scheme.

It follows that as the share of population enrolled in HMOs grew in the late 1980s a larger fraction of the population was subject to practices favored by HMOs and, in the case of deliveries it is natural to expect that this would lead to an observed decline in the cesarean rate. An early indication of this possibility was given in Tussing and Wojotowycz (1992) who reported that the 1986 cesarean rate in a sample of New York HMOs was 13 percent smaller than that for a sample of women enrolled in FFS practice. However, there is an immediate problem with such a direct comparison since omitted health characteristics which may be correlated with enrolling in an HMO may simultaneously influence the probability of delivering by cesarean section.[4]

[4] The selection issue arises because healthier patients, perhaps with unobservable tastes for preventive care, are reported to enroll more frequently in managed care than in traditional insurance; see Cutler and Reber (1988). Healthier patients however, are also less likely to undergo a cesarean. Stafford (1990), Kizer (1988), McCloskey,

While cost conservative measures are perhaps the most important determinants of lower cesarean rates in HMOs, other mechanisms were plausibly at work in depressing non-HMO cesarean rates in this time period as well. An argument in Phelps (1986) suggests that the growth in ob/gyn supply since the 1970s may have produced changes in the medical practice, raised competition and reduced possibilities to induce demand. A related mechanism is the competition between HMO and FFS providers for the same pool of consumers. Specifically, increased enrollments in HMOs could force FFS physicians to mimic HMO practices in order to contract with HMOs and tap into the growing pool of HMO enrollees. In fact, there is indirect evidence for this hypothesis. A 1996 ACOG report reveals that HMOs became an important source of customers for FFS ob/gyns, with 80.1 percent of all FFS ob/gyns participating in salaried managed care positions in 1994, up from 67.2 percent in 1991, even though FFS ob/gyns earned up to 25 percent higher incomes in 1994.

HMOs might also contribute to decreasing the cesarean rate in non-HMOs by encouraging the diffusion of practices favored by managed care. This line of argument draws from physician learning models, e.g., Phelps (1992), which suggest that physicians tend to follow the dominant behavior of other physicians in their local markets, especially when there is professional uncertainty about the "correct" treatment. One explanation consistent with this observed phenomenon is that court decisions typically rest on "local patterns of practice" as the norm in judging malpractice or negligence cases. If such herding indeed takes place, expansions in HMOs over time would make the practice patterns of HMO physicians dominant, encourage non-HMO physicians to embrace the practices favored by HMOs and thereby reduce the cesarean rate.

These mechanisms provide a framework in which evidence of a negative statistical relation between HMO growth and cesarean rates may be given a causal interpretation, which is what we test for in the empirical section. There are two obvious problems the causality framework faces. First, it is possible that while HMO market shares may have some true causal effect in decreasing cesarean rates, HMO market shares may be endogenously determined by unobservables that simultaneously influence cesarean rates. A compelling argument may be that HMOs appear to locate predominantly in up-county, wealthier areas where unobservable tastes for preventive care are higher; but, conditional on risk factors, observable health characteristics and demographics, cesarean rates may also be systematically correlated with such tastes; see Dranove and Wehler (1994). The regression analysis will control for the possibility of such correlated unobservables by using a set of seven state laws that regulate HMO operations, to provide instrumental vari-

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Petitti and Hobel (1992), Tussing and Wojotowycz (1992, 1994) are examples of some studies modeling the determinants of cesareans at the individual level without taking account of the endogenous determination of insurance choice.

ables estimates of causal effect of HMO market shares on cesarean rates. We will consider the plausibility of the exogeneity of state laws below.

A second concern is that there is no true causal mechanism that links HMO market shares with cesarean rates, and it is unobservable contemporaneous changes in tastes for natural deliveries rather than the change in health insurance markets that led to the observed decline in cesareans. Under this scenario the regression analysis will reflect spurious correlations between HMO market shares and cesareans. I will argue against this hypothesis because to my knowledge there is no survey or other evidence of such a discrete change in tastes that could independently account for the large decline in cesareans after 1989 (although it is possible tastes account for some of the reduction). If the hypothesis is true the regression analysis will not be able to distinguish between spurious correlation and a causal mechanism for HMO activity. Therefore, a more reasonable hypothesis for the paper is the following: there are mechanisms that predict a negative causal relation between the growth in HMOs and a decline in cesarean rates. Empirical evidence of a negative relation will be consistent with a causal role for HMO market shares.

3 Theory and Related Research

3.1 A Model of Induced Demand, Market Shares and Cesarean Rates

Let an ob/gyn have a utility function of the form

$$U = U(Y, L, I), \quad U_Y, U_L > 0,\, U_I < 0$$

where Y represents physician income, L is leisure and I is some normalized level of inducement. $U_I < 0$ reflects the assumption physicians derive disutility from knowingly inducing demand, due perhaps to reputation or ethical considerations. This formulation is related to the previous work of Newhouse (1970), Gruber and Owings (1996) and Dranove (1988), extending it to accommodate competition among health insurers.

Let m denote managed care and t correspond to traditional insurance by which we will refer to ob/gyns paid by conventional indemnity insurance. We will use $\bar{B}$ to denote the stock of births in the market, and i to represent the probability that an ob/gyn will induce demand for a cesarean for a given birth or, the "inducement per birth" as in Gruber and Owings (1996); total inducement, I, is simply $\bar{B}i$. Denote $\zeta(i)$ as the cesarean delivery rate, with $\zeta'(\cdot) > 0$ and $\zeta''(\cdot) = 0$.[5]

Let ϕ represent the market share of managed care firms (i.e., the fraction of the local population that is enrolled in managed care). Then

[5] Since some fraction of births are appropriately diagnosed as requiring cesarean delivery, we assume $\zeta(0) > 0$.

$$B^m = \phi \bar{B}, \quad B^t = (1-\phi)\bar{B}$$

and for $j = m, t$

$$C^j = \zeta(i) B^j, \qquad N^j = (1-\zeta(i))B^j$$

where N represents the total number of natural deliveries and C denotes the total number of cesarean deliveries. Let $\overline{L}$ represent the stock of leisure, τ_n the average time per unit spent on a natural delivery and τ_c the average time per unit on a cesarean delivery with $\Delta\tau = \tau_n - \tau_c > 0$. The symbols Y_n and Y_c will respectively denote the ob/gyn reimbursements for a natural and cesarean delivery respectively with $\Delta Y = Y_c - Y_n > 0$. It follows that leisure and total physician payment for child delivery are

$$Y^j = Y_n N^j + Y_c C^j, \qquad L^j = \overline{L} - \tau_n N^j - \tau_c C^j. \tag{3.1}$$

Assume utility to be additive. Subject to the constrains in (3.1) ob/gyn of type j ($j = m, t$) chooses i to solve

$$\max_i U^j(Y^j, L^j, I^j),$$

yielding the first order condition

$$U_Y^j(\cdot)\ [\zeta'(i)\Delta Y] + U_L^j(\cdot)[\zeta'(i)(\Delta\tau)] + U_I^j(\cdot) = 0.$$

That is, the ob/gyn chooses i^* to balance the marginal gains in utility from income and leisure with the marginal disutility from inducement. Noting that $\Delta Y = 0$ for managed care, we see the immediate implication of capitated payments: at the optimum i^*, there is no income-driven inducement in managed care. Instead at the optimum, in managed care i^* solves

$$U_L^m(\cdot)[\zeta'(i)\Delta\tau] + U_I^m(\cdot) = 0.$$

To facilitate comparisons assume equal market shares. It then follows by U_Y,$U_L > 0$ and $U_I < 0$, that

$$i^{m*} < i^{*t} \Longrightarrow \zeta(i^{m*}) < \zeta(i^{t*})$$

or, the optimal cesarean rate is strictly lower for managed care. Define $\bar{C}$ as the total number of cesarean births and let $\bar{c}$ denote the weighted average of the cesarean rates in managed care and traditional insurance. Then,

$$\bar{c} = \zeta(i^m)\phi + \zeta(i^t)(1-\phi)$$

and the comparative statics of interest is given by:

$$\frac{d\bar{c}}{d\phi} = \zeta'(i^t)\frac{di^t}{d\phi} + \phi\left(\zeta'(i^m)\frac{di^m}{d\phi} - \zeta'(i^t)\frac{di^t}{d\phi}\right) + (\zeta(i^m) - \zeta(i^t)) \tag{3.2}$$

which cannot be signed a priori. As managed share grows (ϕ increases), the income pressure on non-managed care ob/gyns increases, so $\frac{di^t}{d\phi}$ is plausibly positive and the first term is plausibly negative. The appendix shows that $\frac{di^m}{d\phi} - \frac{di^t}{d\phi} < 0$, implying that the sign of the second term will depend on $(\zeta'(i^m) - \zeta'(i^t))$, which cannot be definitively signed.

The third component of this equation is the difference in the cesarean rate between managed care and traditional insurance, $\zeta(i^m) - \zeta(i^t)$. Central to this paper is the hypothesis that due to the absence of pecuniary incentives, threat of dismissal and utilization reviews, cesarean rates are likely to be lower in managed care relative to traditional insurance. However, this is a question for empirical analysis. Also, we cannot omit the possibility that $\zeta(i^m) - \zeta(i^t) > 0$ for some range of ϕ. For instance, when ϕ is very low and a very small fraction of births are covered by managed care it is plausible that managed care ob/gyns offset the decline in utility from lower total income with more utility from leisure; i.e., by greater inducement. Therefore equation (3.2) cannot be a priori signed.

Note that holding market shares constant the response of cesarean rates to changes in fertility is unambiguously negative as

$$\frac{d\bar{c}}{d\bar{B}} = (1-\phi)\frac{di^m}{d\bar{B}} + \phi\frac{di^t}{di\bar{B}} < 0 \tag{3.3}$$

which is the basis for previous studies' examination of changes in cesarean delivery rates to declining fertility rates.[6] Without controlling for contemporaneous changes in managed care market shares, however, there is no a priori reason to expect the cumulative effect of declining fertility and increasing managed care market shares to lead to a reduction in cesarean rates; instead, as equation (3.2) shows this is indeterminate and a question for empirical analysis.

3.2 Related Research

A number of micro-level studies have examined variation in cesarean rates by payor source. Early research includes Keppel et al (1980) and Goldfarb (1984), who found that self-paying patients had the lowest rates, patients covered by private insurance had the highest rates, and speculate that financial incentives underlie these differences. Managed care was much less prevalent in

[6]The derivation follows since both $\frac{di^m}{dB}$ and $\frac{di^t}{diB} < 0$ as shown in the Appendix.

the time-frame of these studies, so there is no direct evidence on differences in cesarean rates between managed care and traditional insurance in this body of work.

More recent research has explicitly examined cesarean rate discrepancy in various forms of managed care (e.g., HMO, IPA, etc.) versus commercial insurance or self pay. These include Stafford's (1990) study in the California hospital setting, Tussing and Wojotowycz's (1994) study of New York city hospitals, and McCloskey, Petitti and Hobel's (1992) study on a particular symptom of cesarean (dystocia); see also Kizer (1988). These studies each find a lower probability of cesarean delivery under managed care versus other forms of insurance in regression analysis that controls for maternal demographics, hospital characteristics and various symptoms of cesareans.

However, each of these studies from a fundamental identification problem. Individual decisions to purchase managed care insurance versus purchasing indemnity insurance (or choosing self pay) may be systematically correlated with unobservable tastes or unobservable health characteristics that simultaneously affect the cesarean probability. In fact, a number of studies have found that individuals who joined HMOs were in some respect "healthier" than those that chose conventional insurance; see for e.g., Cutler and Reber (1998), Glied (2000), Cutler and Zeckhauser (2000). However, healthier women may be less likely to undergo cesareans. For instance, unobservable tastes for risky behavior (e.g., disposition for smoking) may raise the likelihood of fetal or maternal distress that are strong determinants of cesareans. Relatedly, numerous studies attribute the higher cesarean rates amongst Hispanic and Black women to the risk factors or poorer health characteristics associated with their lower socioeconomic status; see, e.g., Placek and Taffel (1990) and Tussing and Wojotowycz (1992). While we will control for a number of observable determinants of health, if there is systematic correlation between unobservables that affect both the likelihood of cesareans as well as HMO enrollments, then variations in HMO market shares cannot identify changes in cesarean rates.

This type of self selection has been long noted in the literature examining cesarean probabilities by payor source (see, e.g., Kizer (1988), Stafford (1990), McCloskey et al (1992), Tussing and Wojotowycz (1994)), but has gone unaddressed in empirical analysis. In their well known work Gruber and Owings (1996) use exogenous changes in fertility to identify the induced demand for cesarean rates; however, differences in managed care versus traditional indemnity is not studied in that work. Barring random assignment of women to insurance type, identification of the causal role of HMO market shares will depend on exclusion restrictions that affect the growth of HMO market shares but do not independently affect cesarean rates. Some studies have suggested using the average number of workers in the area as a predictor of HMO growth that may not independently affect health outcomes (see Baker and Browning (1999) for an example from the breast cancer literature; Mitchell and Cromwell (1986) for a similar instrument). However, as suggested in Dranove and Wehler (1994), such instruments may be correlated with both tastes for surgery and underlying determinants of health that in turn affect cesarean probabilities.

Therefore, in this paper the identifying variation will be obtained from a number of state laws that foster or discourage HMO entry, operation and growth. There are at least 24 well defined state regulatory laws that govern the operation of HMOs (HCFA, 1996). Some of these, e.g., capital or reserve requirements, enumeration of basic benefits, prohibitions against deceptive advertising, provision of grievance mechanisms and utilization reviews are uniform across all states and not a source of variation. Of the remaining, there are seven regulations that are plausibly important in HMO locating and expansion decisions, that vary quite widely across states. These include laws that mandate consumer representatives on HMO boards, policymaking roles for subscribers, annual relicensing to operate HMOs and external peer review, each of which plausibly discourage HMO entry or growth; and, laws that might be perceived as encouraging HMO growth, e.g., mandating public and private employers (typically with a minimum number of employees) to offer an HMO option, and accommodation of the Federal HMO law which overrides specific state laws that limit the operations of HMOs.

This source of identifying variation will be sufficient to estimate the causal effect of HMO market share on cesarean rates if unobserved variables that affect the growth in HMOs do not themselves influence the passage of laws or their stringency. Concerns that state laws governing the operation of HMOs emerged as a response to the size or past behavior of HMOs will be alleviated by noting that most of the laws used in this paper were enacted in the late 1970s and early 1980s,[7] substantially preceding the 1991-1992 time period studied in this paper and substantially preceding the dramatic growth in managed care activity. A more thorough discussion of the exclusion restrictions and their regional distributions is given in the next section. In addition the regression analysis will give statistical over-identification tests to empirically verify the exogeneity assumptions.

4 Data and Empirical Analysis

The study is facilitated by a 1989 federal law that requires every birthing center to record the mode of delivery on the birth certificate. These data come from the NCHS which also documents detailed parental, hospital and other birth characteristics of each of the more than 4 million babies born in 1991-1992 in the United States. These data are merged with HMO

[7] One exception is the state dual choice requirement which was introduced in 1973, but underwent substantial amendments in 1985.

market share data from the AAHP for 1991-1992, years in which 99.7 percent of babies were born in a birthing center.

We focus on this time period for two reasons. First, the post-1992 market share data available to this paper were very noisy. Second, as noted before, a variety of structures (e.g., IPAs, PPOs, POS) emerged after the early 1990s blurring the distinction between managed care and traditional indemnity while "HMO" in the early 1990s was a fairly homogeneous and well defined organization. The HMO and NCHS data are merged with county characteristics from the County Business Patterns (CBP), state health regulatory laws from the Health Care Financing Administration (HCFA) and health care facilities from the Area Resource Files (ARF).

The most important drawback of the data is that payor source cannot be identified by birth. Instead, we have measures of HMO enrollment by zip code which are aggregated by counties to compute HMO market shares, measuring the fraction of a local market's (insured) population that is enrolled in the HMO. Counties are used to approximate local markets, since deliveries are not a specialized service for which patients travel far for treatment.[8] HMO data collected by the AAHP are reported only for those zip codes with population size greater than 100,000; this pares the data set to a sample of approximately 400 counties per year. The empirical strategy is to use HMO county market shares rate as indicative of the likelihood that a delivery in that county is covered by an HMO. In keeping with the aggregation of the HMO variable, individual level data from the NCHS are aggregated up to the county level and converted to rates using population data from the NCHS and ARF. With the exception of state regulatory laws all data are at the county level.

Summary statistics are presented in Table 2a. Some striking patterns of HMO activity and cesarean rates emerge from studying the data: nationwide southern states have the lowest HMO market shares and the highest cesarean rates, while western states have the highest HMO activity and the lowest cesarean rates. Of course causality cannot be immediately inferred from these sample statistics since the southern states also have disproportionately larger fractions of women with higher predispositions for cesareans (for example black women, women on AFDC and women in poverty).[9] In the data although aggregate cesarean rates fell between 1991 and 1992, there was substantial underlying heterogeneity across counties. In particular cesarean

[8]Some researchers choose to use Health Care Service Areas (HCSAs), which are larger areas such as groups of counties, to approximate the relevant market (e.g., Makuc et al (1991)). However, HCSAs are not precisely defined, and further, while HCSAs may be superior measures for those services which are less commonly available in every county (such as MRI scans), they may be less relevant for deliveries.

[9]Black women are poorer on average and are less likely to be insured (US Census Bureau 1996), limiting their access to prenatal care, raising the likelihood of poorer maternal health, and thereby raising the likelihood of a cesarean (NCHS, 1994).

rates also rose in approximately 100 counties, although the mean increase in these counties was only .016 percentage points.

Analogously, although the mean HMO market share rose between 1991 and 1992 there is substantial variation in managed care activity across counties. Several local markets exhibit a zero percent HMO market share in both years, and HMO activity actually decreased in 74 counties between 1991 and 1992. Part of the unequal distribution in HMO activity emerges because HMOs appear to prefer locating in wealthier, predominantly white-collar regions with well developed medical infrastructures where their clientele base will be larger; see Baker and Brown (1999). In the sample studied here, the mean HMO market share in the above-average-income markets is 20.1 percent, while it is only 11.7 percent in markets with below-average income.

As described above, this paper uses variations in seven state laws that regulate the growth and performance of managed care institutions to estimate the casual effects of HMO growth on cesarean rates. Each of these variables is an indicator variable taking the value 1 if a particular law is in place (in 1991-1992), and zero otherwise. Descriptive information of these exclusion variables is given in Table 2b.

There are some broad patterns that we can draw from the distribution of laws. On average it appears that the regulatory environment in the northeastern states is conducive to managed care activity (e.g., by mandating Dual Choice, accommodation of Federal HMO, and providing the HMO alternative), discouraging laws that inhibit the growth in managed care (e.g., annual relicensing, external reviews); the regulatory environment in southern states may be perceived as being the least conducive to managed care growth. On average, it appears that the western and northcentral states lie between these regions in the stringency of laws regulating managed care. However, as shown in Table 2b, these are only broad patterns. For example, northeastern states are also likely to require a policy role for subscribers and require a consumer representative on board which could both be perceived as enforcing higher standards of care and may affect HMO entry decisions on the margin; these laws are less prevalent in southern states. It is possible, however, that these latter quality assurance measures are less important in HMO locating decisions relative to some of the other regulatory measures listed in Table 2b.

What drove the differences in the adoption or stringency of these regulatory laws across the states? There is little theoretical or empirical research on this question. Glied (2000) argues that despite an early hostile regulatory environment, federal and state laws in the early 1980s may have been spurred by the dramatic growth in health care costs and the methods in which managed care sought to address to many of the market failures in health insurance: e.g., moral hazard on the part of both consumers and physicians, and asymmetric information between patient and physician. One might speculate that differences in regulatory history, consumer activism and recognition of the possible benefits of managed care led to variations in the regulations that evolved across regions. We will not explore the causes behind the escalation of managed care activity or its variations across regions, but take them as exogenous to the growth in HMO market shares. Critical to the empirical strategy is our belief that the laws that affect managed care growth and enrollments are uncorrelated with the unobservables that affect the disposition for cesareans.

4.1 Empirical Methodology

There are two principal issues that empirical analysis will address. First is the plausible endogeneity of the measure of HMO market shares. Second, as the comparative statics results of Section 3 suggest a possibly nonlinear effect on HMO market shares we will estimate both linear and nonlinear regression models. To this end, the causal effect of the growth in HMO market shares on cesarean rates will be empirically estimated by variants of the model:

$$CRate_{it} = \alpha + f(HMO_{it}) + Fert_{it}\beta_1 + x_{it}'\beta_2 + z_{it}'\beta_3 + \mu_i'\beta_4 + \varepsilon_{it} \tag{4.1}$$

where $CRate$ is the cesarean section rate in county i and year t, HMO is the HMO market share, $Fert$ is the county fertility rate, x is a vector of demographic characteristics including variables relating to race, age, education, income, number of hospital beds per capita, z is a vector of severity variables including previous cesarean, birthweight, fetal distress, maternal distress and μ consists of a full set of state fixed effects.

We will consider two specifications for the function $f(HMO_{it})$, given by[10]

$$\begin{aligned} \text{Specification 1} \quad &: \quad f(HMO_{it}) = HMO_{it}\delta_1 \\ \text{Specification 2} \quad &: \quad f(HMO_{it}) = HMO_{it}\delta_1 + HMO_{it}^2\delta_2. \end{aligned}$$

If HMO market shares were exogenously determined, then estimation of the model in equation (4.1) would lead to correct causal inference for the effect of HMO growth on cesarean rates. However, as described above we are concerned with the endogenous determination of HMO market shares. To address this issue we use instrumental variables (IV) methods to identify and estimate the causal effect of HMO market shares on cesarean rates. For the linear model in

[10] We have also estimated other flexible estimation methods including a semiparametric method in Newey, Powell and Vella (1999); those results are a part of the original version of this paper entitled "Organization of Health Care Markets and the Delivery Choice: Semiparametric Estimates" which may be requested from the author. The more flexible methods yield some differences in point estimates at the tails and are much less precisely estimated due to the high degree of multicollinearity present in such semiparametric regression models.

Specification 1, we will use two stage least squares (2SLS), while for Specification 2 we will use nonlinear IV (NLIV) methods. In both cases, we will generate the instruments by estimation of the reduced form:

$$HMO_{it} = c + Laws_i'\pi_0 + Fert_{it}\pi_1 + x_{it}'\pi_2 + z_{it}'\pi_3 + \mu_i'\pi_4 + v_{it} \tag{4.2}$$

where *Laws* is the vector of seven state laws described in Table 2b. For NLIV, we will use the predicted value $\widehat{HMO}_{it}$ and its square $\widehat{HMO}_{it}^2$ as the instruments. As is well known from Hausman (1983), in the nonlinear case, 2SLS estimates using the predicted values of the variables in place in the endogenous variables leads to biased and inconsistent estimates of the parameters. Therefore, we will replace 2SLS with NLIV methods for the nonlinear models.

The remaining variables in equation (4.1) are characteristics that are expected to explain some of the variation in cesarean rates. Following Gruber and Owings (1996), exogenous decreases in the fertility rate should increase the income pressure on ob/gyns and raise the cesarean rate. Numerous studies have found that specific characteristics such as older mothers, very young mothers, Black, less educated, unmarried (e.g., McCloskey et al (1992), Tussing and Wojotowycz (1992, 1994)), raise the probability of a cesarean, so we control for each of these features. All regressions control for the per capita income of the county, and symptoms of cesareans by including the previous cesarean rate, low and high birthweight rate. In addition, some regressions control for female labor supply, fetal distress rate, the breech presentation rate and the maternal distress rate. We also control for size of the county hospital by using the average number of beds per capita, and control for economic shocks across the two years with a year dummy. To mitigate the possibility that the results reflect fixed differences across states, we also use a full set of state fixed effects.

4.2 Results

Table 3 reports ordinary least squares (OLS) and IV results from estimation of equation (4.1), also giving the reduced form results. We first discuss the linear results. In column (1), consistent with our priors we find a small negative relation between HMO market share and the cesarean rates. However, our primary concern is that these estimates are biased due to the possible endogeneity of HMO market shares. Correcting for the endogeneity via two-stage least squares in column (2) increases the HMO coefficient in absolute value almost tenfold to -.126, which is statistically different from zero. Inflation of the HMO coefficient is suggestive of unobservables that are positively correlated with both cesarean rates and HMO market shares. Our first set of results indicate that a .1 percentage point increase in HMO market shares is associated with a .0126 percentage point decrease in cesarean rates. At 1991 cesarean rates, this is equivalent to a 5.4 percent reduction in cesarean rates.

We find in column (2) that our controls for the percentage of deliveries in a county that are high birthweight (weighing more than 4500 grams at birth), low birthweight (weighing less than 2500 grams at birth) and previous delivery each bear the expected positive sign and are each significant. Also, consistent with Gruber and Owings' hypothesis, the coefficient on fertility is negative and significant supporting the supposition that, conditional on managed care growth, decreases in fertility led to induced demand for cesareans. Relative to the left-out category (Age < 20) the cesarean rates are statistically equivalent for mothers between 20-35 years, but higher for mothers that are 35 years or older. Relative to Hispanics (the left-out group), as the percentage of Black mothers increases so does the cesarean rate, but there is no statistical difference for White mothers.

Many of the instrumental variables in column (4) bear the expected sign, and are statistically significant at the .05 error level. For instance, the presence of dual choice and accommodation of the Federal HMO law in a state are positively correlated with HMO market shares, while mandating policy roles for subscribers and open enrollment have a negative correlation with HMO market shares. Mandating employers to provide an HMO alternative is of an unexpected sign, but insignificant. Contrary to our priors, laws that require a consumer representative on HMO boards has a positive coefficient and is significant. One explanation for this result is that this variable may not itself represent a substantive barrier to managed care growth, and may be absorbing the effects of other unobservable factors that are conducive to HMO operations.[11]

The reduced form has a fairly high adjusted R-squared, and the F-test rejects the joint insignificance of all parameter estimates at the .01 error level. To determine the validity of the instrumental variables, a test of over-identification was implemented for the TSLS regression. The $\chi^2(6)$ statistic of the reduced form regression in column (4) is estimated to be 17.90. Thus, the test of over-identifying restriction rejects the null hypothesis of the joint invalidity of the exclusion restrictions used in estimation, at the .05 percent error level.

The results in column (2) support the basic contention of this paper: by severing the link between income and mode of delivery, encouraging preventive care and promoting cost conservative measures, the growth in HMO market shares results in reducing the national cesarean rates. The empirical results find that the strong, negative and statistically significant relationship between the percent of insured population in HMOs and cesarean rates persists despite controls for hospital size, region, income, demographics and risk factors associated with higher

[11] For instance, it is apparent from Table 2b that states which are more likely to mandate consumer representation are also the states whose regulatory environment is otherwise most conducive for managed care growth.

cesarean rates. Before we consider alternative hypotheses that could produce these results, we examine the robustness of the estimates to a change in specification.

Column (3) gives nonlinear IV estimates of equation (4.1) using a quadratic in HMO market shares. Although the choice of a quadratic seems somewhat arbitrary, we have found that the inclusion of higher powers leads to substantively small differences in the results.[12] Using the reduced form results from column (4), we now use the predicted value of HMO as well as its square as instrumental variables. In column (3) we find that although the NLIV estimate of the linear term is barely significant at the .1 error level, the quadratic is highly significant. The point estimates on the remaining variables are fairly unchanged in value and precision. Evaluating at the sample market share mean of .154, the implied NLIV effect for a .1 percentage point increase in HMO market shares is a .0108 percentage point reduction in cesarean rates. Comparing this result with that obtained in column (2) indicates that ignoring possible nonlinearities somewhat overstates the true HMO effect on cesareans.[13] In a test for examining the validity of the over-identifying restrictions it was found that the $\chi^2(5)$ statistic for the NLIV regression was slightly lower than for the TSLS regression; it was estimated to be 16.70. This rejects the null hypothesis of the non-exogeneity of the joint invalidity of the exclusion restrictions used for the NLIV regression at the .05 error level.

What do the estimates imply? Between 1991 and 1992 the national cesarean rate fell .05 percentage points from .221 to .216, while average HMO market shares rose .083 percentage points. Holding other variables fixed and using the conservative estimate in Table 3, this implies that the growth in managed care activity can explain only .0149 (or 29.8 percent) of the contemporaneous reduction in cesarean rates. The number of cesarean births declined by 16455 between 1991 and 1992 (CDC, 1998). Using HIAA (1991) estimates that the average cesarean cost \$3106 more than the average natural childbirth in 1991, the net savings from births that were delivered naturally in lieu of a cesarean is approximately \$51 million; of this approximately \$15 million can be attributed to the growth in managed care. Although this numeric is a small fraction of total health care costs, it must be noted that this estimate does not take into account other costs such as the longer post-delivery hospital stays and neo-natal intensive care that are associated with cesarean deliveries (Marieskind, 1979).

[12] We have also implemented semilinear and semiparametric methods which are found to yield small differences in the results and a greater amount of imprecision, especially in the tails of the data.

[13] We can also use the NLIV estimates to evaluate varying marginal effects at different market share levels. For example, the implied marginal effect on cesarean rates for a .1 percentage point increase in market shares is .0174, when evaluated at the western average market share. However, the standard errors are correspondingly higher when the estimates are evaluted at this level.

4.3 Alternative Models

We next consider whether the obtained results are logically consistent with alternate explanations than the ones we have offered, examine their robustness to alternate specifications and examine the effects of HMO market shares on outcomes other than the aggregate cesarean rate.

An immediate concern is that regions of low managed care market share (and therefore high traditional indemnity insurance market shares) are regions in which ob/gyns are more likely to diagnose risk factors such as maternal distress, fetal distress and dystocia that are strong predictors of cesarean rates; see e.g. Tussing and Wojotowycz (1992). This is a standard induced-demand hypothesis in which ob/gyns who have pecuniary incentives for cesarean delivery justify their choice by subjectively labelling births as being high-risk. Evidence that ob/gyns may be engaging in this practice is given in McCloskey, Petitti and Hobel (1992), and evidence of this type of excessive coding more generally in the Medicare program is given in Carter, Newhouse and Relles (1990).

While the empirical results have controlled for a variety of risk factors such as age, previous cesarean, low birthweight, high birthweight and race, we address this potential problem by including additional controls for the percent of mothers in county i and year t whose deliveries were labelled as maternal distress, fetal distress or breech presentation. Results from re-estimating the linear two-stage least squares and nonlinear IV estimates are given in Table 4, omitting results for all but the coefficients of interest.

We find that these additional controls do not qualitatively change the results we obtained previously. In the linear model the coefficient on HMO in column (1) remains statistically significant, and the point-estimate is almost unchanged from Table 3. We find that the NLIV-implied estimates in column (2) attenuate relative to the previous results. Specifically, for a .1 percentage point increase in HMO market shares the NLIV results now indicate a .096 percentage point decline in cesarean rates. While all three risk factors bear the positive coefficient, in both columns the coefficient on fetal distress is statistically insignificant at conventional levels. Although this result implies that an increase in the riskiness of births via fetal distress has no effect on cesarean rates, we suggest an alternate interpretation: it is very likely that fetal distress is strongly correlated with other risk factors such as low birthweight, high birthweight and possibly maternal distress, so that once these other symptoms have been controlled for it is possible there remains no statistically discernible independent effect of fetal distress.[14]

[14] Previous studies have found fetal distress to have a positive coefficient on the cesarean probability despite controls for maternal distress, low birthweight and breech; e.g., Tussing and Wojotowycz (1992) and Gruber and Owings (1996). These studies use data on individuals. It is possible that the aggregation of individual data to county rates removes a lot of variation that leads to the result found here.

A second and related concern is that our results could be reflecting regional variations in demands for cesareans arising from variations in female labor force attachment. This will affect our estimates if labor force attachment rates are systematically correlated with the level of managed care activity, and independently affect cesarean probabilities. While there is some evidence that labor force participation is correlated with the reduction in fertility (Olsen (1994)), there is no clear a priori correlation between tastes for cesareans and labor force attachment: although cesareans may reduce the uncertainty in the timing of deliveries, the average post-operation recovery period for cesareans is longer; Placek and Taffel (1980). Similarly, one may speculate that labor force attachment rates are positively correlated with managed care activity (e.g., HMOs may infer a large pool of prospective consumers when the county labor participation rate is larger), although there is little empirical research on this question.[15] In either of these cases, ignoring variations in labor force attachment will lead to omitted variables bias in our estimates.

We use data on state labor force attachment from the Area Resource Files (ARF) to examine this hypothesis and report the results in Table 5. Inclusion of the labor force participation rate attenuates the HMO effect. We find that in the linear specification of column (1) the coefficient on HMO market share declines in absolute value to -.099, but is less precisely estimated than the previous regression models. The NLIV estimates similarly decline, implying a .088 percentage point reduction in cesarean rates for a .1 point rise in market shares. The coefficient on labor force participation itself is negative and highly significant implying that conditional on income, demographics and other risk factors, the incidence of cesareans is declining in labor force participation rates. The inclusion of the labor force variable has no large effects on most other variables, except on fertility whose coefficients attenuate significantly in both columns (1) and (2).[16]

It is instructive to consider whether this result is a function of other omitted characteristics that are correlated with labor force participation. Suppose women in the labor force are more likely to be insured than women out of the labor force. Then, because the uninsured are generally of poorer health (Kasper, 1986), one would expect the labor force participation rate to have a negative correlation with conditions symptomatic of poorer health, e.g., maternal and fetal distress. This would overstate the true causal effect of labor force participation rates on the cesarean rate because such risk factors are excluded from columns (1)-(2) Alternatively, labor force attachment (and delayed childbearing) may be positively correlated with such risk factors

[15] The sample correlation between HMO market shares and the female labor force attachment is .3007.

[16] This last result is suggestive (but not conclusive) of a positive correlation between fertility and female labor force attachment, which is contrary to our priors.

in which case our results may be understating the true effect. In columns (3)-(4) of Table 5 we append the previous regression with controls for risk factors. Inclusion of these additional controls reduces both the effects of labor force participation, as well as the estimated HMO effect on cesarean rates. The attenuation implies that the cesarean rate reduction associated with a .1 percentage point increase in HMO market shares is now .0094 and .008 in columns (3) and (4) respectively, evaluating NLIV at the sample mean.

In column (5) we test another hypothesis based on an assumption that there is some fraction of deliveries in the population that is correctly diagnosed as requiring cesareans. If managed care is driving down the cesarean rate it is natural to expect that they are doing so by first shifting those cesareans that are purely due to induced-demand rather than those that are medically necessary. One implication is that if past expansions in managed care have already reduced the cesarean rate, further increases in managed care penetration may have progressively smaller effects in doing so. Therefore, if a marginal increase in HMO market share does in fact have a causal role in reducing the cesarean rate, this effect should be greatest where HMO market share is smallest and be substantially reduced where HMO market share is already large.

We test this hypothesis by introducing an interaction between HMO market shares and an indicator for market shares that are greater than the median HMO market share[17]:

$$CRate_{it} = \alpha + HMO_{it}\beta_0 + (Median_{it} * HMO_{it})\beta_1 + Fert_{it}\beta_2 + x'_{it}\beta_3 + z'_{it}\beta_4 + \mu'_i\beta_5 + \varepsilon_{it}$$

where $Median_{it}$ is a binary indicator that takes the value 1 if the market share in county i in year t is greater than the median market share in the data and 0 otherwise.

If this hypothesis is correct, then HMO market share coefficient β_0 should remain negative while β_1 should be positive. Using instruments for both the interaction variable and HMO market share variable,[18] our results in column (5) do not give support to this hypothesis at conventional significance levels: β_1 is positive, but it is statistically significant at less than the 70 percent confidence level. Although any inference based on these estimates will be imprecise we note, with this caveat, that these estimates imply that for a .1 percentage point increase in HMO market shares the generic reduction in the cesarean rate is .0101 percentage points; but for regions with greater than median managed care activity the reduction is slightly less at .009. In results not reported here we have also estimated this model using breakpoints other than the median market shares; this change does not produce results that are more supportive of the hypothesis.

[17] The median HMO market share is .135. Using the median rather than the mean makes the estimates less influenced by outliers.

[18] A separate reduced form is estimated using the interaction as the dependent variable.

A question of interest is whether the HMO effect on the cesarean rate is working through its effect on primary cesareans (cesareans on first births), repeat cesareans, or both.[19] There is substantial empirical evidence that the probability of a cesarean conditional on a previous cesarean delivery is tremendously high; e.g., Marieskind (1979), Tussing and Wojotowycz (1992). Evaluating whether reductions in the aggregate cesarean rate are due to reductions in the primary or repeat cesarean rates can be instructive in understanding the mechanisms through which managed care insurance results in lowering the cesarean rate.

Despite historically high rates, primary cesareans declined appreciably from more than 16 percent in 1989 to 15.1 percent in 1992, while the contemporaneous reduction in repeat cesarean rates was smaller: from 8.5 percent to 8.2 percent (CDC, 1998). The growth in managed care could in principle lead to reductions in both these rates. If a large fraction of primary cesareans were due to induced demand, then the rise in HMO practices and the absence of pecuniary incentives for HMO ob/gyns should directly result in a reduction of the primary cesarean rate. In fact, Gruber and Owings (1996) find that up to 32 percent of primary cesareans may have been the result of induced demand. While ob/gyns have traditionally defended repeat cesareans on medical grounds, recent evidence has shown that VBAC rates in HMOs have risen more sharply than those in traditional insurance; e.g. Stafford (1991), PHC (1995). Thus it is possible that there exists a statistical association between managed care activity and the reduction in each of these rates. To assess this hypothesis we estimate the regression models with the repeat cesarean rate and primary cesarean rate as dependent variables, and give the results in Table 6.[20]

We find that while increases in HMO enrollments have a strong negative association with the primary cesarean rate given in column (1), its relationship with the repeat cesarean rate given in column (3) is statistically indiscernible from zero. In fact, although the F-statistic in column (3) rejects the joint insignificance of all the coefficients at the .05 error level, it is still fairly low. These findings are echoed by the NLIV estimates for repeat cesareans in columns (2) and (4). For all columns, the risk factors are found to be positively associated with both primary and repeat cesareans, but appear to have a stronger effect on primary cesareans. Interestingly, we find that the inducement effect for fertility is substantially higher for primary than for repeat

[19] The cesarean rate is the sum of primary and repeat cesarean rates. Conditional on having a previous cesarean, if a birth is not repeat cesarean then it is a VBAC. The primary (repeat) cesarean rates are measures as the number of primary (repeat) cesareans normalized by the total number of births in a county.

[20] We note here the relation between "Previous Cesarean" and "Repeat Cesarean". At the individual level, all individuals that are classified as having repeat cesareans must have had a previous cesarean (by definition). However, having had a previous cesarean does *not* necessarily imply that the individual has a repeat cesarean; the individual may have a VBAC. Therefore, repeat and previous cesarean are not perfectly collinear.

cesareans. For primary cesareans, our linear and nonlinear results indicate that a .1 percentage point increase in HMO market shares is associated with between .0082 and .009 percentage points reduction in primary cesarean rates. The absence of evidence linking managed care with repeat cesarean rates suggests that perhaps due to the medical uncertainty surrounding the success of VBACs in this time period (see ACOG, 1996), HMOs played a role in reducing the aggregate cesarean rate largely by reducing the number of first-time cesareans.

Finally we consider, as in Gruber and Owings (1996), whether there are asymmetric responses in cesarean rates to increases versus declines in managed care penetration. It is plausible that once managed care has grown and managed care practice patterns have evolved as the "norm", exogenous declines in managed care market shares have little effect on raising the cesarean rates. A candidate explanation for this pattern would be that malpractice laws typically evaluate the appropriateness of a procedure based on local standards of practice; see Phelps (1992). Therefore, if past expansions in managed care have led natural childbirth to be the dominant practice pattern, ob/gyns in traditional insurance may have reduced incentives to induce demand for cesareans due to the threat of malpractice. Consequently, future exogenous declines in managed care penetration may have a negligible effect on raising cesarean rates.

To test this hypothesis we construct an indicator variable ("Decline") which takes the value 1 for the 74 counties in which HMO market shares declined from 1991 to 1992. We then construct the interaction of this variable with HMO market share, and estimate regressions of the form:

$$\begin{aligned} CRate_{it} &= \alpha + HMO_{it}\beta_0 + Decline_{it}\beta_1 + (HMO_i * Decline_{it})\beta_2 + Fert_{it}\beta_3 \\ &\quad + x'_{it}\beta_4 + z'_{it}\beta_5 + \mu'_i\beta_6 + \varepsilon_{it} \ . \end{aligned}$$

The coefficient β_2 is a measure of the differential effect on cesarean rates due to declines in HMO market share. The results are given in Table 7.

If the cesarean rate responds symmetrically to rises and declines in HMO market shares, then β_0 should unchanged from a regression without the *Decline* and interaction variables, and β_2 should be zero. We find no evidence of asymmetric responses in regressions with and without controls for female labor force attachment, as the coefficient β_2 is statistically indistinguishable from zero. However, we find a positive discrete effect on cesarean rates in counties where managed care activity fell ($\beta_1 > 0$).

These results give evidence of symmetric responses in cesarean rates to increases and declines in managed care activity. One explanation for this result is that the growth in managed care until 1992 may have been too recent a phenomenon for its practice patterns to have evolved as the "local standard of care". In this case, in the event of exogenous declines in managed care activity ob/gyns in traditional insurance may continue to respond as before to the fee differential between vaginal and cesarean deliveries. Then, because the observed cesarean rate is the sum of rates in managed care and traditional insurance exogenous decline in managed care will be associated with an observed rise in cesarean rates.

4.4 Quality of births

The previous section examined the short run effects of HMO growth on cesarean rates. This section focuses on whether, conditional on cesarean rates, increased HMO activity impacts the quality of births. Numerous studies have documented differential quality in service between traditional fee-for-service and managed care systems; e.g., Cutler, McClellan and Newhouse (1997), Feldman and Scharfstein (1999), McLaughlin (1988). This section focuses on three specific measures of quality: neo-natal infant mortality rates, birth injury/complications and Apgar scores.

The regression results above have shown that increased HMO activity is correlated with shifting births from cesareans to natural deliveries. Although the absence of pecuniary incentives in managed care may be the primary reason for the decline in cesarean rates, some fraction of this decline may also be attributed to the increased pressure to maintain lower costs in managed care. There is anecdotal evidence, for example, that HMO physicians face substantial pressure (including threat of dismissal) to follow practice patterns established by the HMO and reduce costs by choosing lower-reimbursed alternatives; see Robinson and Steiner (1998).

It is possible therefore that at the margin an HMO ob/gyn may have increased incentives to perform a natural delivery.[21] It could be argued that the higher rates of VBACs in HMOs (Stafford, 1991) is indicative of this phenomenon. However, since performing a natural delivery when a cesarean is medically necessary could result in delivery injuries/complications including infant mortality, HMO market shares may be correlated with an increase in observed delivery injuries/complications or neo-natal infant mortality rates. Alternatively, the NCHS has found that by actively encouraging prenatal care pre-existing conditions may be better detected under managed care, reducing the likelihood of complications at birth. A 1996 CDC study on 1990 births found that fetal mortality was four times higher for mothers who had no exposure to prenatal care,[22] which would indicate a negative correlation between HMO market shares and neo-natal infant mortality.

[21]For example for the marginal delivery if there is some indication but no strong, compelling factors suggesting a cesarean, it is conceivable that an HMO ob/gyn chooses natural delivery.

[22]The study also indicates that access to prenatal care is greatly higher for educated, White mothers. One example of managed care prenatal intervention that could result in lowered cesarean rates is the repositioning of breech positions, prior to the first trimester. Another is the detection of pre-existing conditions.

To test these hypotheses, we estimate instrumental regressions of the form:

$$Q_{it} = \alpha + HMO_{it}\beta_0 + CRate_{it}\beta_1 + Fert_{it}\beta_2 + x'_{it}\beta_3 + z'_{it}\beta_4 + \varepsilon_{it}$$

where Q_{it} is alternatively the birth injury/complication rate, neo-natal infant mortality rate and percentage of low Apgar scores in county i at year t.[23] By focussing on quality measures that are typically observed after the decision to perform a cesarean has been made, we mitigate any concerns for reverse causality.

We use the six hour neo-natal infant mortality rate since this, rather than a 28-day rate, is more likely to be directly related to the delivery. We control for the county cesarean rate because cesareans result in roughly twice the blood loss as natural deliveries and anesthesia is more frequently administered for cesareans, each of which could raise the likelihood of birth injuries/complications (CDC, 1995). Also, injuries from surgical procedures are more commonly reported for cesarean deliveries. As both fetal deaths and Apgar scores are plausibly correlated with the number and degree of birth injuries, we include birth injury rates as a regressor in the infant mortality and Apgar rate regressions. Apgar rate data were available for only 199 counties, so we use three region dummies in place of the state fixed effects for that regression.

Regression results are given in Table 7. Before discussing the results note that in the birth injury/complication and Apgar rate regressions (columns (2)-(3)) although the F-statistics reject the null hypothesis that all the coefficients are jointly zero at the 95 percent level, the very low pseudo R-squared indicates that the chosen regressors explain very little of the dependent variable variation over this time period. While we discuss the results, we caution against placing any strong emphasis on them because of the fairly weak power of the regressors in explaining the dependent variables' variation.

We find that the coefficient on HMO market shares is statistically insignificant in all but the birth injury/complication rate (where it is significant at the .1 error level), suggesting that in the 1991-1992 period the growth in HMOs was not associated with tandem declines in these quality measures. Similarly, once various risk factors have been controlled for the cesarean rate has no statistical effect on either the injury rate or Apgar rate; however it has a small but statistically negative association with the neo-natal infant mortality rate. While there is some informal notion that cesareans are more likely to results in infant deaths (see, e.g., Cox and Schwartz , 1990), this observation could simply be a result of the induced correlation between low birthweight and cesarean rates. That is, low birthweight babies are both more likely to be

[23] We use the Apgar score measured at 5 minutes after birth, which reflect conditions of heart rate, respiration, muscle tone, reflex, irritability and color, and are utilized in predicting a newborn's probability of survival (NCHS, 1994). An Apgar score can range from 1 (poor health) to 10 (good health); the low Apgar score rate is the percentage of babies with scores under 7.

delivered cesarean and more likely to die; see CDC (1998). Thus, once low birthweight has been controlled for it is statistically possible for cesareans to be positively associated with neo-natal infant mortality rates. As expected, risk factors such as low weight, breech, maternal distress and fetal distress have large positive effects on both the birth injury rate and the neo-natal infant mortality rate. In addition, in the neo-natal infant mortality rate regression, the coefficient on birth injury/rate is found to be positive and statistically significant at the .05 error level.

In summary, our results give some support for the hypothesis that the initial growth in managed care is not associated with declines in certain qualitative measures of deliveries in this time period. Since the mid 1990s, the popular press has questioned the propriety of VBACs and raised concerns that the cost pressure on HMOs has led to a rise in birthing injuries through VBACs.[24] More recently, Frigoletto et al (1999) have criticized government initiatives in promoting VBACs and found systematic evidence that by 1995 birth injuries were often tied to vaginal labor in cases where, in previous years, ob/gyns would have opted for a cesarean.[25] Thus, whether the continued rise in managed care over the 1990s led to any change in the patterns found in this section, is an important question for future research.

5 Conclusions

A standard theory of induced demand makes salient predictions about delivery choices when ob/gyns are confronted with an exogenous adverse shock to their income: ob/gyns will shift deliveries away from natural childbirths to the more highly reimbursed alternative, cesarean delivery. Equivalently, the exogenous reduction in fertility in 1989-1996 should have been coincidental with a rise in cesarean rates, but the raw data indicate otherwise. This paper reconciles this apparent contradiction of the induced model by examining the contemporaneous growth in managed care activity. Managed care severs the link between obstetrics' delivery choices and financial reimbursement, eliminating one of the principal incentives in induced demand for cesareans. Thus, the dramatic increase in managed care activity in the early 1990s should exact an effect opposite from that of declining fertility on cesarean rates. This paper examines whether the data are consistent with this hypothesis.

We find no contradiction of the induced demand model. The evidence shows that declining fertility was associated with an increase in the cesarean rate, but that this effect was negated by a reverse effect from growing managed care activity that led to the observed reduction

[24]See, e.g., Korte (1998); and, "Hospital, HMOs shouldn't dictate birthing options", from The Hampton Union, 2001. (Available at www.seacoastonline.com/news/hampton/h10_23_e1.htm)

[25]Besides the obvious risks for patients, these authors pointedly note that reverting to a cesarean after attempting a VBAC is twice as costly than the cesarean may have cost.

in cesarean rates. This relation is estimated precisely despite controls for birth severity, risk factors, demographics, female labor force attachment, hospital size, and fertility. We confront the endogeneity problem that arises from the possibly joint determination of cesarean rates and managed care enrollments by using state variation in laws that promote or hinder managed care growth and operations. Our results indicate that the managed care activity can statistically explain approximately 30 percent of the reduction in cesareans. Simple calculations at 1991-1992 managed care levels and 1992 cesarean rates indicate that the medical (dollar) savings from the HMO-induced reduction in cesarean rates amounted to over \$15 million, as a conservative estimate. We find no reduction in the contemporaneous "quality" of births, but temper our findings due to reduced precision in this estimation.

There is more work for future research. First, there is a question of whether reductions in managed care enrollments after 1991-1992 led to any change in the patterns obtained in this study. After peaking in the early 1990s, managed care enrollments declined in the next few years (AAHP, 1994). There would be a useful extension in understanding whether, for e.g., VBACs emerged as the "practice norm" in obstetrics' choices, preventing a reversal towards the high cesarean rates of the late 1980s upon exogenous reductions in managed care. Second, it is of importance to consider the "quality" aspect of deliveries that this study has begun to consider. Although there is little evidence of reductions in quality in the 1991-1992 time period, recent concerns in both the academic and popular press have suggested that cost pressures in HMOs may have led them to adopt VBACs or natural childbirths for deliveries that might medically be more suitable for cesareans; see Frigoletto et al (1999). These are topics of both academic as well as policy interest, and should be a first priority for future research.

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A Appendix A

The comparative statics of interest is

$$\frac{d\bar{c}}{d\phi} = \zeta'(i^t)\frac{di^t}{d\phi} + \phi \left(\zeta'(i^m)\frac{di^m}{d\phi} - \zeta'(i^t)\frac{di^t}{d\phi}) \right) + (\zeta(i^m) - \zeta(i^t)).$$

The first term is unambiguously negative. The first order conditions for the traditional indemnity insurance and the managed care physician are respectively

$$\begin{aligned} U_Y^t(\cdot)\ [\zeta'(i^t)\Delta Y] + U_L^t(\cdot)[\zeta'(i^t)(\Delta\tau)] + U_I^t(\cdot) &= 0 \\ U_L^m(\cdot)[\zeta'(i^m)\Delta\tau] + U_I^m(\cdot) &= 0. \end{aligned}$$

Totally differentiating, replacing $B^m = \phi\bar{B}$, $B^t = (1-\phi)\bar{B}$ and noting that all terms except U_{II}, U_{LL} and U_{YY} are positive, gives

$$\begin{aligned} \frac{di^t}{d\phi} &= -\frac{U_{YY}^t(\cdot)\zeta(i^t)[-\Delta Y^2]\zeta'(i^t) + U_{LL}^t(\cdot)\zeta(i^t)[-\Delta\tau^2]\zeta'(i^t) - U_{II}^t(\cdot)i^t}{U_{YY}^t(\cdot)\Delta Y^2\zeta'(i^t)^2(1-\phi) + U_{LL}(\cdot)\zeta'(i^t)^2(1-\phi)\Delta\tau^2 + U_{II}^t(\cdot)\phi} > 0 \\ \frac{di^m}{d\phi} &= -\frac{U_{LL}^m(\cdot)\zeta(i^m)\Delta\tau^2\zeta'(i^m) + U_{II}^m(\cdot)i^m}{U_{LL}^m(\cdot)\zeta'(i^m)^2\phi\Delta\tau^2 + U_{II}^m(\cdot)\phi} < 0, \end{aligned}$$

and therefore

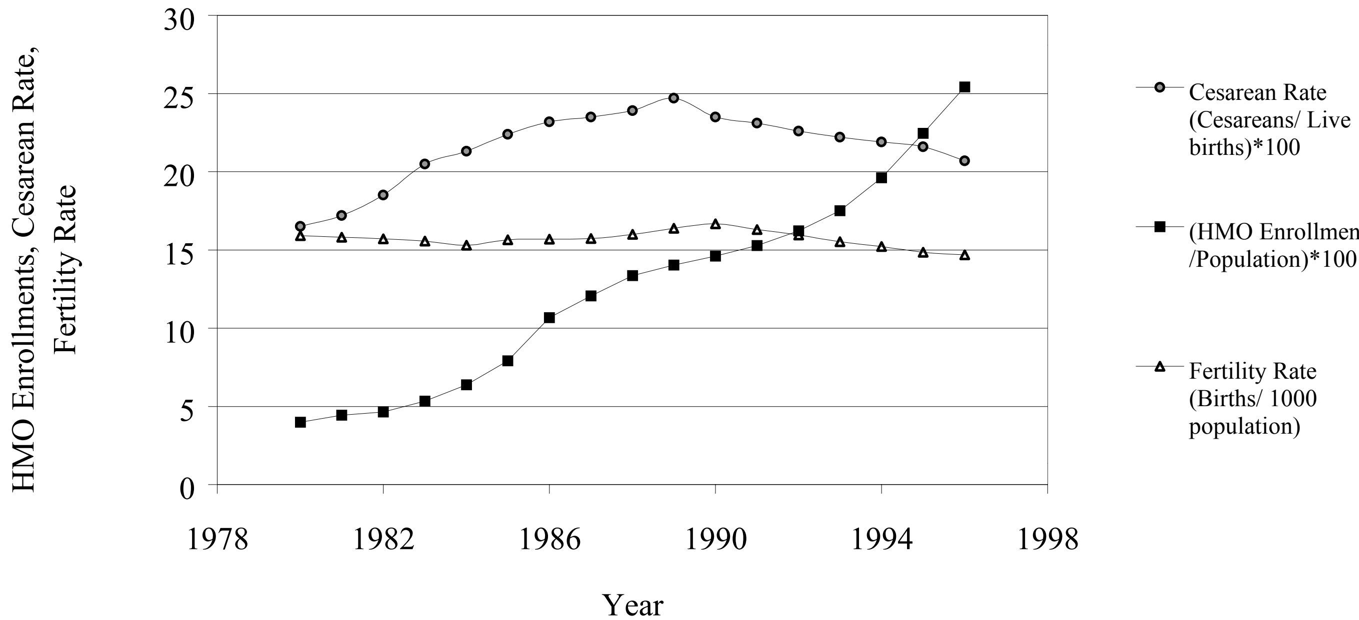
$$\frac{d(i^m - i^t)}{d\phi} < 0.$$

A similar result shows that

$$\begin{aligned} \frac{di^t}{d\bar{B}} &= \frac{U_{YY}^t(\cdot)\zeta'(i^t)\Delta Y[Y^t/\bar{B}] + U_{LL}(\cdot)\zeta'(i^t)\Delta\tau[\{1-\phi\}[-\tau_n(1-\zeta(i^t) - \tau_c\zeta(i^t)] + U_{II}^t(\cdot)(1-\phi)i^t}{U_{YY}^t(\cdot)\Delta Y^2\zeta'(i^t)^2(1-\phi) + U_{LL}(\cdot)\zeta'(i^t)^2(1-\phi)\Delta\tau^2 + U_{II}^t(\cdot)\phi} < 0 \\ \frac{di^m}{d\bar{B}} &= \frac{U_{LL}^m(\cdot)\zeta'(i^m)\Delta\tau[\phi[-\tau_n(1-\zeta(i^m) - \tau_c\zeta(i^m)] + U_{II}^m(\cdot)\phi i^m}{U_{LL}^m(\cdot)\zeta'(i^m)^2\phi\Delta\tau^2 + U_{II}^m(\cdot)\phi} < 0. \end{aligned}$$

We are thus able to sign the first two components of equation (3.2).

Figure 1



HMO Enrollments, Cesarean Rates and Fertility Rates (1980-1996)

Table 1
Scheduling incentives for Cesareans

	Percent of All Deliveries	Percent of Cesareans Deliveries
Sunday	10.60	6.61
Monday	14.30	15.41
Tuesday	16.40	17.64
Wednesday	15.71	17.64
Thursday	15.31	16.91
Friday	15.94	20.60
Saturday	11.71	5.147

Source: Author's calculations based on a random sample of NCHS data for 1991 and 1992. There are 8032 total deliveries in this table and 1363 cesarean deliveries.

Table 2a: Sample Means

	All Counties	West	South	North Central	Northeast
HMO Market share/100	.154	.236	.108	.149	.163
population	[.110]	[.133]	[.089]	[.090]	[.105]
Cesarean Rate	.218	.196	.229	.218	.226
(Cesareans/100 births)	[.052]	[.044]	[.052]	[.060]	[.033]
Fertility Rate	.016	.017	.017	.016	.013
(Births /population)					
Percent of births:					
Low Birthweight	.006	.005	.006	.007	.007
High Birthweight	.034	.033	.036	.034	.034
Fetal Distress	.041	.037	.042	.039	.042
Maternal Distress	.040	.040	.044	.038	.041
Breech Presentation	.039	.039	.038	.040	.039
Neo-natal infant mortality	.021	.019	.017	.026	.019
Birth injury/complication	.002	.002	.002	.003	.002
Low Apgar Score	.020	.0001	.030	.026	.0001
Percent of Mothers:					
Primary Cesarean	.136	.123	.148	.131	.139
Repeat Cesarean	.081	.072	.080	.086	.087
Age < 20	.171	.165	.197	.179	.131
Age 20-30	.572	.565	.567	.585	.565
Age 30-35	.189	.191	.173	.178	.224
Age 36-40	.059	.068	.054	.050	.069
Age 41-45	.007	.009	.006	.005	.008
Age > 41	.0001	.0002	.0001	.0001	.0002
Labor force participation	.215	.206	.216	.215	.220
Black	.140	.041	.249	.118	.104
White	.827	.884	.727	.863	.870
Married	.721	.735	.690	.734	.730
High School or less	.575	.537	.603	.588	.553
College or more	.399	.377	.389	.405	.421
Other					
Hospital beds per capita	4.85	3.61	5.39	4.78	5.11
	[2.66]	[2.35]	[3.16]	[2.42]	[2.12]
Per capita income	19.7	19.63	19.10	18.94	21.91
(1000s of dollars)	[4.18]	[4.17]	[3.95]	[2.88]	[5.26]
n	823	155	243	238	187

Note: Data from NCHS, AAHP and ARF as described in text. Where present, standard deviations in parenthesis. Hispanic is the left out category for race. Apgar rate data are computed from 160 counties.

<u>Table 2b</u>
Distribution of Laws

	Number of states mandating law	Percent of states mandating law, within region			
		West	Northeast	South	Central and N.Central
Consumer Rep.	12	.37	.38	.06	.18
Policy Role for Subscribers	31	.81	.75	.46	.63
Annual Relic.	23	.62	.13	.8	.36
Open Enroll.	17	.37	.38	.2	.45
Review	12	.27	.13	.26	.27
HMO Alt.	11	.27	.5	.2	.22
Dual Choice	14	.27	.63	.2	.36
Federal HMO	21	.45	.67	.33	.18

<u>State Laws Regulating HMO Operations in 1991-1992:</u>

Consumer Rep: 1 if the state mandates consumer representatives on HMO boards, 0 otherwise.
Policy Role: 1 if the state mandates a policymaking role for HMO subscribers, 0 otherwise.
Annual Relic.: 1 if the state mandates an annual relicensure for HMOs, 0 otherwise.[1]
Ext. Review: 1 if the state mandates an external body (whether the State Health. Commisioner, or another regulatory agency) to monitor the quality and enforcement of standards in HMOs, 0 if either the State explicitly forbids such monitoring or has no relevant provision for monitoring.
Open Enroll: 1 if the state mandates open enrollment, 0 otherwise.
HMO Alt: 1 if the state mandates employers to offer an HMO alternative, 0 otherwise
Dual Choice: 1 if the state requires that employers with a minimum number of employees offer an HMO alternative, 0 otherwise.[2]
Federal HMO: 1 if the state mandates accommodation of the Federal HMO law, 0 otherwise.[3]

[1] In 1991, states either had annual relicensing procedures or granted certificates for indefinite operations. One exception was Michigan which used a 3-year licensing period. For the variable *Annual Relicensure*, Michigan takes the value 1.
[2] Where this law is enforced in 1991, the minimum number of employees is 25 in all cases, except the state of Washington which requires 50 employees.
[3] Section 1311(c) of the Federal HMO act makes explicit provisions to counteract state laws that prevent "HMOs from doing business as such and from operating in accordance with section 1301 of the Act." This Section of the act provides more lax conditions to counteract state regulations on stringency of i) medical society approval of services provided by HMOs, ii) number of physicians on government agencies, iii) capital or financial reserves; and, v) state laws that prevent HMOs from complying with the Federal HMO act (see State Law Compliance Guide, Section 1).

Table 3

OLS, Two Stage least squares (TSLS) and NLIV estimates

Dependent variable:	Cesarean Rate			HMO Market share
	(1): OLS	(2): TSLS	(3): NLIV	(4)
HMO Market share	-.035	-.126	.014	
	(.019)	(.061)	(.008)	
(HMO Market share)^2			-.399	
			(.113)	
Fertility Rate	-.929	-.935	-.963	-.369
	(.351)	(.351)	(.349)	(.649)
Low Birthweight	1.79	1.81	1.80	.325
	(.285)	(.285)	(.283)	(.526)
High Birthweight	.978	.961	.961	-.181
	(.084)	(.085)	(.084)	(.157)
Previous Cesarean	.260	.279	.268	-.108
	(.028)	(.031)	(.027)	(.049)
Age 20-30	-.256	-.207	-.247	.829
	(.146)	(.127)	(.147)	(.186)
Age 30-35	-.087	.112	.097	-.190
	(.266)	(.110)	(.153)	(.274)
Age 35-40	-.083	.125	.102	-.671
	(.266)	(.078)	(.093)	(.494)
Age > 40	.354	.229	.224	2.29
	(.087)	(.063)	(.083)	(1.64)
Married	.007	-.037	-.033	-.454
	(.038)	(.051)	(.051)	(.071)
White	.059	.020	.023	-.360
	(.039)	(.049)	(.049)	(.071)
Black	.090	.043	.045	-.396
	(.032)	(.025)	(.025)	(.079)
Hospital Beds/capita	.003	.002	.002	-.008
	(.0008)	(.001)	(.001)	(.001)
Per capita income	.0002	.0005	.0006	.004
	(.0006)	(.0007)	(.0007)	(.001)
Federal HMO Law				.018
				(.010)
Dual Choice				.008
				(.004)
HMO Alternative				-.004
				(.015)
Policy Role for Subscribers				-.022
				(.010)
Open Enroll.				-.053
				(.012)
Review				-.008
				(.005)
Consumer Rep.				.078
				(.018)
Adjusted/Pseudo R^2	.3046	.3560	.3661	.5180
F statistic	[69,754]=7.1	[69,754]=7.0	[70,751]=7.08	[75,748]=14.5
χ^2 statistic				$\chi^2[6] = 17.9$

Notes: Standard errors in parenthesis. For column (2), (3), std. Errors account correctly for the first step. All regressions include state dummies, a year dummy, education variables, with n=823. Hispanic and Age < 20 are left out variables. For TSLS and NLIV, pseudo R^2 is reported. The χ^2 statistic determines the validity of the exclusion restrictions used in estimation of column (2).

Table 4

Cesarean Rates and Risk Factors: Fetal Distress, Maternal Distress, Breech Presentation

Dependent Variable:	Cesarean Rate (1): 2SLS	Cesarean Rate (2): NLIV	HMO Market share (3)
HMO Market share	-.124	.022	
	(.059)	(.013)	
(HMO Market share)^2		-.386	
		(.111)	
Fertility Rate	-1.06	-1.09	-.400
	(.348)	(.347)	.(654)
Low Birthweight	1.65	1.64	-.208
	(.308)	(.306)	(.572)
High Birthweight	.891	.895	-.237
	(.085)	(.085)	(.158)
Previous Cesarean	.202	.203	-.097
	(.027)	(.029)	(.050)
Fetal Distress	.413	.409	-1.89
	(.321)	(.322)	(3.20)
Maternal Distress	.819	.837	.277
	(.162)	(.167)	(.114)
Breech Presentation	.426	.417	.103
	(.085)	(.085)	(.161)
Federal HMO Law			.020
			(.010)
Dual Choice			.012
			(.011)
HMO Alternative			-.006
			(.015)
Policy Role for Subscribers			-.023
			(.010)
Open Enroll.			-.055
			(.012)
Review			-.005
			.009
Consumer Rep.			.079
			(.018)
Adjusted/Pseudo R^2	.3262	.3345	.5204
F-statistic	[72,751]=7.32	[73,750]=7.36	[78, 745]=15.7
χ^2 statistic			χ^2 [9]=18.23

Notes: Two-step standard errors in parenthesis. Both regressions also include the entire set of regressors from Table 3, columns (2), (3). There are 823 observations used in each regression. The χ^2 statistic determines the validity of the exclusion restrictions used in estimation of column (1). A separate over-identifying test was implemented for the regression in column (2).

Table 5

Cesarean Rates and Labor Force Attachment

	(1): TSLS	(2): NLIV	(3):TSLS	(4): NLIV	(5): Median
HMO Market share	-.099	.020	-.094	.022	-.1009
	(.058)	(.011)	(.057)	(.014)	(.073)
(HMO Market share)^2		-.353		-.348	
		(.110)		(.115)	
Median*HMO Market share					.011
					(.009)
Labor force attachment	-.506	-.515	-.482	-.489	-.460
	(.087)	(.086)	(.103)	(.102)	(.104)
Fertility Rate	-.827	-.854	-.967	-.999	-.821
	(.344)	(.342)	(.340)	(.337)	(.358)
Low Birthweight	1.89	1.88	1.64	1.64	1.87
	(.279)	(.277)	(.299)	(.297)	(.286)
High Birthweight	.980	.980	.892	.898	.949
	(.083)	(.083)	(.083)	(.083)	(.085)
Previous Cesarean	.258	.262	.259	.253	.254
	(.030)	(.031)	(.031)	(.030)	(.032)
Fetal Distress			.403	.405	.405
			(.364)	(.366)	(.368)
Maternal Distress			.769	.758	.760
			(.172)	(.179)	(.173)
Breech Presentation			.466	.457	.455
			(.083)	(.083)	(.082)
Pseudo R-squared	.3337	.3439	.3626	.3721	.3631
F-statistic	[70,753]=7.75	[71,752]=7.84	[73,750]=8.31	[74,749]=8.38	[74,749]=8.10

Notes: Two-step standard errors in parenthesis. Both regressions also include the entire set of regressors from Table 3, columns (2), (3). There are 823 observations used in each regression. In results not reported, the reduced form is estimated for both TSLS and NLIV, including Labor force attachment as a regressor.

Table 6

Primary Cesarean Rates and Repeat Cesarean Rates

Dependent Variable:	Primary Cesarean Rate		Repeat Cesarean Rate	
	(1): TSLS	(2): NLIV	(3): TSLS	(2): NLIV
HMO Market share	-.094	.013	-.079	.084
	(.046)	(.009)	(.080)	(.116)
(HMO Market share)^2		-.330		-.193
		(.074)		(.151)
Fertility Rate	-1.12	-1.07	-.487	-.469
	(.471)	(.472)	(.218)	(.228)
Low Birthweight	.712	.762	.695	.602
	(.194)	(.186)	(.418)	(.418)
High Birthweight	.523	.647	.509	.549
	(.053)	(.055)	(.115)	(.116)
Previous Cesarean	.258	.262	.458	.456
	(.030)	(.031)	(.130)	(.134)
Fetal Distress	.223	.208	.253	.258
	(.338)	(.326)	(.452)	(.464)
Maternal Distress	.746	.813	.625	.645
	(.318)	(.303)	(.343)	(.301)
Breech Presentation	.500	.515	.527	.511
	(.053)	(.052)	(.114)	(.115)
Pseudo R-squared	.3648	.3784	.1769	.1797
F-statistic	[72,751]=8.65	[71,750]=8.74	[72,751]=3.85	[71,750]=3.77

Notes: Two-step standard errors in parenthesis. Both regressions also include the entire set of regressors from Table 3, columns (2), (3). There are 823 observations used in each regression. Instruments for HMO variables are estimated from reduced form regressions not reported here.

Table 7

Testing for Differential Responses to Increases and Declines in HMO Market shares

	(1): TSLS	(2): TSLS
HMO Market share	-.132	-.103
	(.075)	(.079)
Decline*HMO	.196	.202
Market share	(.332)	(.382)
Decline	.077	.069
	(.047)	(.043)
Fertility	-1.43	-1.10
	(.351)	(.370)
Labor force attachment		-.460
		(.103)
Low Birthweight	1.84	1.53
	(.344)	(.355)
High Birthweight	.620	.813
	(.101)	(.115)
Previous Cesarean	.212	.209
	(.024)	(.031)
Fetal Distress	.449	.450
	(.333)	(.337)
Maternal Distress	.736	.740
	(.166)	(.184)
Breech Presentation	.278	.433
	(.095)	(.110)
Pseudo R-squared	.3634	.3685
F-statistic	[76, 747]=8.00	[77,746]=8.05

Notes: Two-step standard errors in parenthesis. Both regressions also include the entire set of regressors from Table 3, columns (2), (3). There are 823 observations used in each regression. Separate reduced forms were estimated for HMO market share, Decline and the interaction variable. Also, for each of these, separate reduced forms were estimated with and without Labor force Attachment as a regressor for column (1) and (2) respectively.

Table 8

Effects on Quality Measures: Two Stage Least squares (TSLS) Estimates

Dependent Variable:	Neo-natal infant mortality rate (1): TSLS	Birth Injury/ Complication Rate (2): TSLS	Low Apgar Score Rate (3): TSLS
HMO Market share	-.013	-.018	.262
	(.016)	(.011)	(.349)
Cesarean Rate	-.083	-.019	-.104
	(.041)	(.005)	(.202)
Labor Force Attachment	.101	.025	-.859
	(.027)	(.015)	(.621)
Fertility Rate	-.069	.028	-2.09
	(.103)	(.051)	(1.84)
Low Birthweight	.277	.029	1.47
	(.093)	(.019)	(1.09)
High Birthweight	.023	.003	.196
	(.025)	(.013)	(.514)
Previous Cesarean	.016	.012	.118
	(.007)	(.004)	(.144)
Fetal Distress	.582	.454	.923
	(.504)	(.250)	(.878)
Maternal Distress	.061	.027	.598
	(.018)	(.009)	(.508)
Breech Presentation	.108	.034	-.002
	(.025)	(.012)	(.670)
Birth Injury/	.313		1.903
Complication Rate	(.078)		(2.63)
Pseudo R-squared	.1849	.0458	.0135
F-statistic	[75,78] = 3.82	[74,79] = 1.97	[28, 348] = 1.04

Notes: In column (3), there are only 320 observations. In the remaining regressions, n=823. Two-step standard errors in parenthesis. Both regressions also include the entire set of regressors from Table 3, columns (2), (3). The reduced form for columns (1) and (3) consisted of birth injury as a regressor; the reduced form for column (2) was estimated without this additional control.