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Physician financial incentives and cesarean delivery:
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The cesarean section is one of the most common surgical procedures, and one of those most studied by economists, for good reasons: it absorbs billions of dollars of health care resources annually, is used with widely varying frequency across regions and across providers, and is potentially responsive to a variety of economic forces, including source of payment, malpractice liability, and financial incentives.

Yet there are few studies of the last of these, financial incentives. One of the most recent and most expansive, by Gruber, Kim, and Mayzlin (1999, hereafter GKM), relates cesarean use to the fee premium paid to physicians by Medicaid when a cesarean delivery is performed instead of a vaginal delivery. In a panel analysis of publicly available childbirth microdata covering nine states for the years 1988-1992, GKM conclude that $434 additional reimbursement (in 1989 dollars), which was the difference in the cesarean fee premium between Medicaid and private insurance in 1989, increases the probability of cesarean delivery by 3.04 to 5.51 percentage points, which was most of the 5.7 percentage point private-Medicaid cesarean differential in 1989. Thus the widely-documented difference in cesarean rates between public and private payers is primarily accounted for by financial incentives. While a few other studies, mentioned below, conclude otherwise, none has received the attention of this one, which has 31 citations, in published and unpublished manuscripts, in Google Scholar as of June 2008.

The accessible data, simple design, and striking conclusions of this study make it a suitable candidate for replication. Accordingly, we obtained the same data used by GKM originally and attempted to reproduce their results, generously assisted by the original authors, who shared their data assembly programs with us. In addition, as standard empirical practice in panel studies of this type has evolved somewhat since the original study was published, we also explored the effect of some of these changes on the findings–particularly the inclusion of state trends, which are clearly observable in the data, in the regression specification. These trends and, to a lesser degree, data assembly issues, are both important. We ultimately conclude that financial incentives may influence cesarean utilization in these data, but, if so, the effect is much smaller than originally advertised.

1. Description of the original study

Since 1988, the Nationwide Inpatient Sample of the Healthcare Cost and Utilization Project (HCUP) has recorded clinical and nonclinical information on roughly one-fifth of all inpatient stays in community hospitals, about one thousand hospitals in various states nationwide in each year. GKM analyzed the 1988-1992 HCUP data for nine states: California, Colorado, Florida, Illinois, Iowa, New Jersey, Pennsylvania, Washington, and Wisconsin. Arizona and Massachusetts, also represented in these data, were omitted because the former had an unusual Medicaid program and the latter had "inconsistent" data on physician fees. This left 42 state*year cells (certain years are missing in two states) containing 365,942 Medicaid-financed deliveries used for analysis.

This analysis related the probability of a cesarean delivery to the Medicaid physician cesarean fee premium (cesarean fee - vaginal fee) in that state in that year, FEEDIFF; state and year fixed effects, μ and τ; and controls for maternal demographics, hospital characteristics, and clinical indicators for cesarean section, X. Estimated using a logit model, the specification is as follows:

$$P(C_{i,s,t}=1) = \Lambda(\beta X_i + \gamma FEEDIFF_{s,t} + \mu_s + \tau_t) \qquad (1)$$

where i, s, and t index individuals, states, and time; C is a dummy that equals one if individual i's delivery was performed by cesarean section and zero otherwise; Λ is the logistic distribution function; and β and γ are regression coefficients. We are mostly interested in the value of γ.

The effect of fee changes was predicted using the marginal effect of a one unit change in FEEDIFF at the mean of the independent variables. The model was estimated with a single clinical control (for a prior cesarean section) and with an expanded set of four clinical controls (adding dummies for breech presentation, fetal distress, and maternal distress). These two models were also estimated with an alternative FEEDIFF measure, the fee premium divided by the average private reimbursement for a vaginal delivery in that state in that year. The results were reasonably similar across all four estimations: as noted above, the predicted effect of a $434 increase in the fee differential ranged from 3.04 to 5.51 percentage points.

2. Replication

We obtained the identical data from the Agency for Healthcare Research and Quality, attempted to recreate the estimation sample using GKM's simple selection criteria (all Medicaid-financed deliveries in the above-listed states), and recreated the original variables. These were taken verbatim from the data except for the clinical indicators, created from standard ICD-9 diagnosis codes, and the cesarean indicator, determined from the Diagnostic Related Group (DRG) code. The FEEDIFF variable, merged in at the state*year level, is provided by GKM in the original study.

We first encounter difficulty reproducing the sample, whose size differs by a factor of three: 947,664 observations to 365,942 observations. This difference accrues partly to the method used to identify the childbirth cases in the HCUP data. We use DRG codes, which are universally reported here. GKM, in contrast, used supplemental ICD-9 diagnosis codes (V270-V277) that are not reported as frequently, as documented in Table 1, which shows the total number of observations accruing under different selection criteria. When weighted by the discharge weight to the universe of U.S. community hospital discharges, the full set of DRG-selected childbirth observations maps closely to total births in the U.S. Vital Statistics, as it should (the small difference reflects out-of-hospital deliveries and multiple births). Unweighted, GKM's ICD-selected sample size is consistently two-thirds of ours, and double that reported in the original (which we cannot account for). Accordingly, we proceed with the larger, DRG-selected sample.

Variable means, presented in Table 2, suggest the two samples are different, but not markedly so. GKM's means include the states of Arizona and Massachusetts, which are excluded from their empirical analysis, so we present means that include, and exclude, these two states. Both ways the cesarean rate is two percentage points higher in the DRG-selected sample, while mothers are slightly older and somewhat wealthier; teaching hospitals are more frequent, and public hospitals less frequent. There is no strong geographic or racial difference across all three samples. Note also that all samples are dominated by just two states, California and Florida, which account for over 60% of the observations used in our regressions. Estimation results, therefore, may be sensitive to the weighting these states receive, or, equivalently, the assumed error structure.

Next we re-estimate GKM's two specifications with the original FEEDIFF measure: Model (1) includes only a prior cesarean dummy as a clinical control, while Model (3) adds dummies for breech birth, fetal distress, and maternal distress. This last variable, not precisely defined, was intended to reflect "complications...arising from either conditions existing prior to pregnancy (i.e., diabetes, hypertension, or infectious diseases) or pathological conditions which develop during pregnancy (i.e., eclampsia or placenta previa)" (pp. 487-488). While these complications are not uncommon, in GKM's means this dummy is one only 0.01% of the time. Their programs show that this dummy was set to one only for maternal distress "not elsewhere classified" (ICD-9 code 669.0), excluding common sources of maternal distress, which have their own identifiers. Accordingly, we run two Model (3) regressions: one that omits the maternal distress dummy, which should be virtually identical to GKM's Model (3) given the rarity of their maternal distress indicator, and one that includes a more robust indicator based on GKM's description above, including diabetes, hypertension, pre-eclampsia, placenta previa, herpes, and maternal distress not elsewhere classified.

The results are in Table 3. Coefficients on the controls rarely differ in sign from those in the original, but frequently differ in magnitude. The standard errors also differ, partly because of the increased sample size and partly because GKM's standard errors are adjusted for state*year clustering, which we will discuss shortly. The coefficient estimates on FEEDIFF are 20% to 30% lower than in the original. These coefficients imply that a $434 fee increase raises cesarean rates by 2-2.5 percentage points.

For this calculation we use the average marginal effect, the theoretically appropriate measure, increasingly recognized as preferable to the marginal effect evaluated at the mean of the independent variables, the measure used by GKM (Fernandez-Val, 2007). The two differ substantially here, as Table 2 shows, because cesarean probabilities are largely bifurcated across mothers: they are high for mothers with clinical indications for a cesarean, and low for those without such indications. Either way, the marginal effect of a fee increase is modest, as is, then, the average of these marginal effects. In contrast, the average of the independent variables falls closer to the middle of the logit distribution, where the marginal effects are relatively large. Consequently, our implied coefficient effects are 40% smaller than in the original, despite a smaller divergence in the coefficient estimates.

3. Estimation and specification

To explore estimation issues, it is instructive to first illustrate the association between fees and cesarean probabilities at the state level, where the "natural experiments" driving this study are conducted. Figure 1 presents three variables for all state*year cells in our estimation sample: FEEDIFF, the unadjusted cesarean rate, and a proxy for the cesarean rate that is adjusted for hospital characteristics, maternal demographics, and prior cesarean section, as in Model (1). This proxy is created by regressing the probability of cesarean delivery on these factors and 42 dummies representing all state*year cells, $m_{s,t}$, as follows:

$$P(C_{i,s,t}=1) = \Lambda(\beta X_i + m_{s,t}) \qquad (2)$$

The coefficients m, which are not directly interpretable but are virtually proportional to adjusted cesarean rates, are graphed in Figure 1. For the most part, adjusted rates closely track unadjusted rates, because the independent variables change little within states over time. Deviations can be traced back to changes in the means of the prior cesarean dummy and total hospital discharges.

Examination of FEEDIFF within states across years reveals virtually no change in four states, modest changes (< \$150) in three, and large changes (> \$150) in two: California and Colorado. These changes are typically monotonic. Similarly, there is little trend in adjusted or unadjusted cesarean rates in some states, and a steady downward trend in others. These trends sometimes coincide with fee changes, raising the spectre of bias.

This is true for California in particular, which combines a dramatic reduction in fees with the largest reduction in adjusted cesarean rates observed in the sample. This is the only *positive* association between fees and cesarean use in any state, and it alone accounts for the econometric estimates reported previously. When the sample excludes California, estimates of the FEEDIFF coefficient, in Table 4, are negative. The inferences drawn from these data depend crucially on whether one accounts for state-specific trends and whether one treats California as an outlier.

Further examination indicates we should do the former but not the latter. The years 1988-1992 represented a turning point in U.S. cesarean rates: the first decline after two decades of increase (Clarke and Taffel, 1995). California led this trend, both in alacrity and vigor: the cesarean rate crested there sooner, in early 1987, and (adjusted for risk) dropped faster, as Figure 1 shows. Stafford, Sullivan, and Gardner (1993, p. 1301) suggest three possible reasons for this: 1) "the well-publicized publication of several articles critical of high cesarean section rates," 2) "increasing public awareness of cesarean section practices," and 3) "changing reimbursement policies of insurers, especially the state Medicaid program," Medi-Cal, which aggressively lowered the cesarean fee premium (in stages) from $563 in October 1986 (AGI, 1987) to $0 in November 1989 (ACOG, 1991). Accordingly, cesarean rates fell for all payers after 1987, with Medi-Cal leading the way: a drop of 3.5 percentage points through 1990, at least twice that of other payers Kaiser, selfpay, and private insurance (Stafford, Sullivan, and Gardner, 1993). Our data show that adjusted cesarean rates continued to drop through 1992.

In summary, reduced reimbursement was an important strategy used to cut cesarean utilization in California in the late 1980s, but was probably not the sole relevant factor; the decline in cesarean rates is roughly, but not closely, coincident with the decline in the fee differential. Under these circumstances it seems most reasonable to include California in the sample and, following recent practice, to include state-specific trends in the specification. These regressions, in the fourth row of Table 4, obtain coefficient estimates on FEEDIFF that are about half those obtained in the trendless estimates (significant in Model 1, not in Model 3).

An alternative, more technical approach arrives at the same conclusion. In the regressions above observations are "clustered" by state*year cells if there are unobserved, time-varying, state-specific factors that also influence the cesarean rate–that is, state*year random effects. Then ordinary least squares produces biased standard errors and inefficient coefficient estimates. The bias and the inefficiency can each be large when a few states dominate the data and when there are serially correlated errors, conditions that are met here. GKM's standard errors are corrected for clustering; we also correct the coefficient estimates, using a generalized least squares regression of the m coefficients in equation (2) on FEEDIFF and the fixed effects:[1]

$$m_{s,t} = \gamma FEEDIFF_{s,t} + \mu_s + \tau_t + \epsilon_{s,t} + \nu_{s,t}, \qquad \epsilon_{s,t} = \rho\epsilon_{s,t-1} + \eta_{s,t} \tag{3}$$

where ϵ is the random effect, which may be serially correlated, and $\nu_{s,t}$ is the sampling error in estimating $m_{s,t}$. The results, in Table 4, are similar to those in the regressions with state trends, accompanied by larger standard errors that sometimes render the coefficient estimate insignificant.

Across the bottom nine regressions in Table 4, which account for state*year clustering, serial dependence, or both, the coefficient estimates center around 0.02. This is about one-third the size of the 0.05-0.07 estimates reported by GKM. Average marginal effects calculated from these estimates imply that a \$100 increase in fees, in contemporaneous dollars, raises cesarean rates by a little more than one-quarter of a percentage point, or that the \$434 nationwide difference in cesarean reimbursement rates between Medicaid and private insurance in 1989 raised cesarean rates by roughly one percentage point. This is one-fourth the size of GKM's original marginal effects evaluated at the mean of the independent variables (and is less precisely estimated).

GKM also present estimations (Model 2 and Model 4) with an alternative measure of financial incentives, the Medicaid cesarean fee premium normalized by the private charge for vaginal delivery. We have re-estimated the first four specifications in Table 3 with this alternative measure. As with the original fee measure, our coefficients are somewhat smaller than in GKM; as in GKM, these estimates imply somewhat larger effects than the original fee measure does. We do not present these estimates, however, because we are unable to reproduce GKM's alternative fee measure where we have the necessary data,[2] and lack the data to recreate it entirely.

4. Conclusion

In summary, we find financial incentives have, at best, a small effect on delivery methods at the margin. In current dollars, a $1000 increase in the reimbursement for performing a cesarean section would increase cesarean delivery rates by little more than one percentage point. This finding jibes with that of the closest related study, by Keeler and Fok (1996), of a cesarean/vaginal fee equalization at one insurer. It contrasts with that of Currie and Gruber (2001), who find even larger effects than GKM using a fully national dataset but a less direct research design.

Our findings imply that financial incentives are not an effective mechanism with which to influence physicians' choice of delivery method. This is consistent with numerous contemporaneous anecdotal reports of unsuccessful attempts by Blue Cross units to lower cesarean rates through reduced financial incentives (Stafford, 1990; Darby, 1992; Ricci et al. 1993).

Our findings also imply that other factors account for the large private payer/public payer gap in U.S. cesarean rates. Analyzing a nationwide sample of births in 1988 and all Florida hospital births in 1992, Grant (2005) finds that clinical risk factors and matching between privately insured mothers and physicians who are predisposed to perform cesareans explains all but one percentage point of this gap. This remaining difference corresponds to the contribution of financial incentives estimated here.

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Table 1. Sample Sizes and Weights in the HCUP Data.

Sample Restrictions	Raw Numbers of Observations: ICD-9 Codes Identify Births	Raw Numbers of Observations: DRG Codes Identify Births	Previous Column Weighted by Discharge Weight	Data from Alternative Sources
Total Births 1988-1992, U.S. Vital Statistics	-----	-----	-----	20,284,061
All Birthing Cases in 1988-1992 HCUP	2,292,373	3,434,491	19,283,280	-----
Above with Primary Payer = Medicaid	731,883	1,028,828	5,797,859	-----
Above with AZ and MA Removed	689,175	947,746	5,396,710	-----
Above with Missing Age Data Removed	see note	947,664	5,396,295	-----
Actual Number of Observations Reported in GKM	-----	-----	-----	365,942

Note: Missing age data appears to be included in the omitted age category in GKM's regressions.

Table 2. Variable Means.

Variable	Original	Replication: All	Replication: Estimation Sample
Number of Observations	365,942 **	1,028,735	947,664
Cesarean Delivery Rate	0.178	0.197	0.197
Clinical Factors: Previous Cesarean	0.093	0.100	0.100
Fetal Distress	0.094	0.103	0.101
Breech Presentation	0.029	0.031	0.031
Maternal Distress	0.0001	0.065 ***	0.065 ***
Median Income: < $15,000	0.053	0.053	0.055
$15,000-$20,000	0.173	0.156	0.158
$20,000-$25,000	0.257	0.247	0.246
$25,000-$30,000	0.207	0.212	0.212
$30,000-$35,000	0.118	0.132	0.130
$35,000-$40,000	0.065	0.072	0.070
$40,000-$45,000	0.026	0.034	0.033
> $45,000	0.021	0.025	0.026
Income Missing	0.080	0.069	0.071
Age: < 20	0.246	0.248	0.248
20-25	0.371	0.339	0.338
25-30	0.222	0.197	0.197
30-35	0.111	0.093	0.094
35-40	not listed	0.031	0.031
40-45	not listed	0.005	0.005
> 45	not listed	0.000	0.000

Race: White	0.201	0.203	0.193
Black	0.085	0.088	0.088
Other	0.252	0.229	0.236
Hospital Characteristics: Teaching	0.343	0.419	0.413
Private for Profit	0.053	0.049	0.052
Public	0.327	0.295	0.311
Number of Discharges (1000s)	22.250	21.547	21.751
State: Arizona	0.021	0.033	----
California	0.331	0.296	0.321
Colorado	0.022	0.019	0.020
Florida	0.283	0.269	0.293
Iowa	0.051	0.040	0.043
Illinois	0.099	0.112	0.121
Massachusetts	0.037	0.046	----
New Jersey	0.025	0.052	0.056
Pennsylvania	0.044	0.046	0.050
Washington	0.027	0.031	0.033
Wisconsin	0.059	0.057	0.062

** The number of observations listed for the original study refers to the estimation sample. The means presented in that study include data from the states Arizona and Massachusetts, which were omitted for estimation. Including those states, the original study means would come from approximately 388,500 observations (365,942)/(1-0.021-0.037). The number of observations in the "Replication: All" column has missing age data removed, and so is slightly smaller than that in the corresponding cell of Table 1.

*** A broader definition of maternal distress is used in the replication, as discussed in the text.

Table 3. Logit Model Estimates (coefficient estimates, with standard errors in parentheses).

Specification → Variable	Original: Model (1)	Replication: Model (1)	Original: Model (3)	Replication: Model (3), no maternal distress	Replication: Model (3), w/ maternal distress
Fee Differential ($100s)	0.0494 (0.0123)	0.0355 (0.0047)	0.0654 (0.0150)	0.0530 (0.0051)	0.0573 (0.0052)
Clinical Factors: Previous Cesarean	2.973 (0.054)	2.780 (0.008)	3.360 (0.061)	3.131 (0.009)	3.220 (0.009)
Fetal Distress	----	----	1.821 (0.039)	1.746 (0.008)	1.775 (0.008)
Breech Presentation	----	----	1.035 (0.338)	3.636 (0.016)	3.724 (0.016)
Maternal Distress	----	----	3.720 (0.045)	----	1.519 (0.010)
Median Income: < $15,000	-0.115 (0.047)	-0.179 (0.021)	-0.069 (0.051)	-0.169 (0.023)	-0.174 (0.023)
$15,000-$20,000	-0.046 (0.035)	-0.109 (0.018)	-0.037 (0.035)	-0.109 (0.020)	-0.103 (0.020)
$20,000-$25,000	-0.019 (0.029)	-0.084 (0.017)	0.000 (0.036)	-0.079 (0.019)	-0.082 (0.019)
$25,000-$30,000	0.042 (0.029)	-0.003 (0.018)	0.047 (0.030)	0.007 (0.019)	0.019 (0.019)
$30,000-$35,000	0.050 (0.033)	-0.022 (0.018)	0.072 (0.032)	-0.004 (0.020)	0.007 (0.020)
$35,000-$40,000	0.062 (0.031)	0.003 (0.020)	0.092 (0.028)	0.012 (0.022)	0.018 (0.022)
$40,000-$45,000	0.082 (0.034)	0.004 (0.023)	0.063 (0.038)	-0.016 (0.025)	-0.009 (0.025)
> $45,000	0.058 (0.049)	0.029 (0.024)	0.079 (0.053)	0.040 (0.026)	0.046 (0.027)
Age: 20-25	-0.035 (0.013)	-0.036 (0.007)	-0.011 (0.015)	-0.024 (0.008)	-0.004 (0.008)
25-30	0.022 (0.017)	0.049 (0.008)	0.038 (0.018)	0.043 (0.009)	0.054 (0.009)

30-35	0.115 (0.021)	0.202 (0.010)	0.101 (0.023)	0.170 (0.011)	0.141 (0.012)
35-40	0.353 (0.030)	0.424 (0.016)	0.335 (0.030)	0.389 (0.017)	0.293 (0.018)
40-45	0.559 (0.044)	0.663 (0.036)	0.587 (0.054)	0.641 (0.039)	0.475 (0.040)
45+	0.546 (0.179)	0.198 (0.149)	0.674 (0.186)	0.269 (0.162)	0.213 (0.164)
Race: White	-0.041 (0.028)	0.026 (0.013)	-0.003 (0.030)	0.039 (0.014)	0.026 (0.014)
Black	-0.109 (0.050)	-0.050 (0.015)	-0.050 (0.060)	-0.012 (0.016)	-0.033 (0.016)
Other	-0.155 (0.040)	-0.112 (0.014)	-0.071 (0.039)	-0.029 (0.016)	-0.014 (0.016)
Hospital Characteristics: Teaching	-0.243 (0.053)	-0.070 (0.007)	-0.439 (0.083)	-0.184 (0.008)	-0.196 (0.008)
Private for Profit	0.081 (0.046)	0.126 (0.013)	0.129 (0.055)	0.155 (0.014)	0.155 (0.014)
Public	0.102 (0.042)	0.043 (0.007)	0.122 (0.057)	0.038 (0.008)	0.025 (0.008)
Number of Discharges ***	0.000 (0.000)	-0.009 (0.000)	0.000 (0.000)	-0.010 (0.000)	-0.013 (0.000)
State and Year Dummies	yes	yes	yes	yes	yes
Interpreting the Fee Coefficient: Marginal Effect at the Mean * 4.34 (percentage points)	3.13 **	2.44	4.14 **	3.64	3.93
Average Marginal Effect * 4.34 (percentage points)	----	1.95	----	2.49	2.62

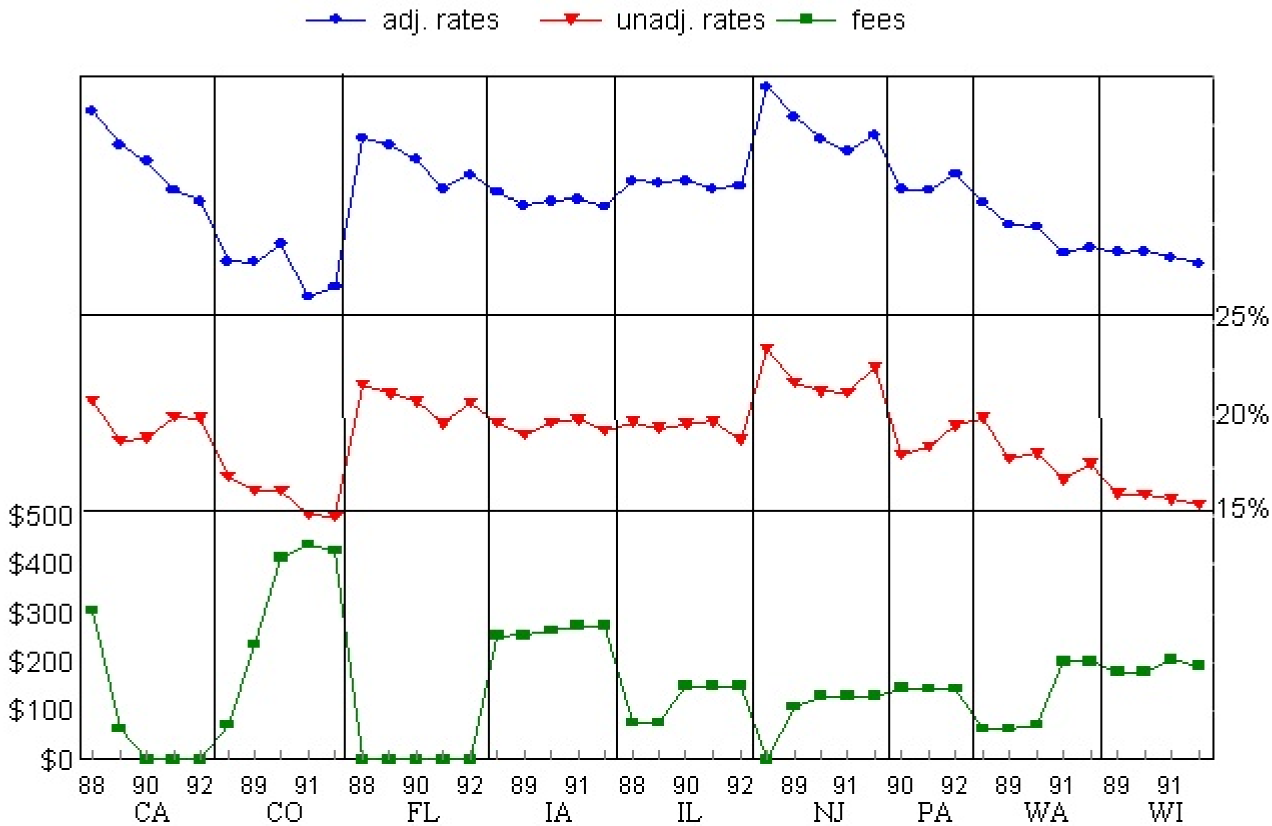
Note: The dependent variable is a binary indicator for cesarean delivery. N=365,942 in the original study and 947,664 in the replication. Standard errors in the original are adjusted for state*year clustering. ** The entries for the marginal effect at the mean in the original study are this author's calculations using the findings presented in the original. The original authors' calculations differ slightly. *** Discharges in the replication, but not the original, are measured in thousands.

Table 4. Alternative Estimates of the Fee Effect (coefficient estimates, with standard errors in parentheses and average marginal effects in brackets, when applicable).

	Model (1)	Model (3), no maternal distress	Model (3), with maternal distress
Original	0.0494* (0.0123)	0.0654* (0.0150)	0.0654* (0.0150)
Replication (Table 2)	0.0355* (0.0047) [0.45]	0.0530* (0.0051) [0.57]	0.0573* (0.0052) [0.60]
Omit California	-0.0176 (0.0138)	-0.0089 (0.0149)	-0.0064 (0.0151)
Add Quarterly State Trends	0.0208* (0.0079) [0.26]	0.0098 (0.0085) [0.11]	0.0092 (0.0086) [0.10]
Two-Step GLS (no autocorrelation)	0.0229* (0.0109) [0.29]	0.0363* (0.0141) [0.39]	0.0341* (0.0127) [0.36]
Two-Step GLS (AR1 random effect)	0.0179 (0.0131) [0.23]	0.0234 (0.0154) [0.25]	0.0250 (0.0146) [0.26]

Note: The dependent variable is literally (in the first four rows) or effectively (in the last two rows–see the discussion in the text) a binary indicator for cesarean delivery. Fees are measured in hundreds of dollars, and average marginal effects are multiplied by one hundred for ease of interpretation. Thus, the average marginal effect gives the expected effect on cesarean rates, in percentage points, of a $100 increase in fees (in contemporaneous dollars). Year dummies are also included in the state trends specification; the coefficient estimates are slightly larger if these are removed. N= 947,664 in the full micro sample; N = 643,182 in the omit California subsample; there are 42 state*year cells in the two-step GLS estimations. * = coefficient significant at $p < 0.05$.

Figure 1. Cesarean Rates and Fees in the HCUP Data, 1988-1992.



Note: Unadjusted cesarean rates are in percentage points. Adjusted cesarean rates are logit coefficients on a full set of state*year dummies, in arbitrary units, in a logit also including all other control variables included in Model (1) described in the text. True adjusted cesarean rates would be approximately linearly related to these coefficients. The X-axis indexes years within states (labeled by their postal code).

References

Alan Guttmacher Institute, 1987. The Financing of Maternity Care in the United States. Alan Guttmacher Institute, New York.

American College of Obstetricians and Gynecologists, 1991. Improving Access to Maternity Care: Obstetric Provider Participation in Medicaid. The American College of Obstetricians and Gynecologists, Washington, D.C.

Clarke, Sally, and Selma Taffel, 1995. Changes in cesarean delivery in the United States, 1988 and 1993. Birth 22 (2), 63-67.

Currie, Janet, and Jonathan Gruber, 2001. Public health insurance and medical treatment: the equalizing impact of the medicaid expansions. Journal of Public Economics 82, 63-89.

Darby, M, 1992. Reimbursement Has Small Impact on C-Section Rates. In: Report on Medical Guidelines and Outcome Research. Capitol Publications, Alexandria, Va.

Fernandez-Val, Ivan, 2007. Fixed effects estimation of structural parameters and marginal effects in panel probit models. Unpublished working paper, Boston University, Boston.

Grant, Darren, 2005. Explaining source of payment differences in U.S. cesarean rates: Why do privately insured mothers receive more cesareans than mothers who are not privately insured? Health Care Management Science 8 (1), 5-17.

Gruber, Jonathan, John Kim, and Dina Mayzlin, 1999. Physician fees and procedure intensity: the case of cesarean delivery. Journal of Health Economics 18 (4), 473-490.

Keeler, Emmett, and T. Fok, 1996. Equalizing physician fees had little effect on cesarean rates. Medical Care Research and Review 53 (4), 465-71.

Ricci, Joseph, L.M. Beam, J.V. Gigliotti, G.F. Rosetti, and T. Wickizer, 1993. An analysis of cesarean deliveries among Pennsylvania Blue Shield subscribers. Pennsylvania Medicine 96 (9), 26-29.

Schwartz, Anne, David Colby, and Anne L. Reisinger, 1991. Variation in medicaid physician fees. Health Affairs 10 (1), 131-139.

Stafford, Randall, S.D. Sullivan, and L.B. Gardner, 1993. Trends in cesarean section use in California, 1983 to 1990. American Journal of Obstetrics and Gynecology 168 (4), 1297-1302.

Stafford, Randall, 1990. Alternative strategies for controlling rising cesarean section rates. JAMA 263 (5), 683-687.

Endnotes

1. This estimator is a slight generalization of a common two-step procedure for dealing with the clustering problem. The generalization is that estimation of equation (3) accounts not just for sampling error in estimating $m_{s,t}$ in equation (2), but also for the covariance in these estimates across time within states (both of which are taken from the covariance matrix estimated in equation (2)). In practice, estimation is simplified by sweeping out state fixed effects in equation (2) before estimating equation (3). Further details are available from the author.

2. This measure is the simple ratio of the "Medicaid fee difference" in GKM's Table 1b (some of whose elements have been confirmed in the original data sources) and the "private payers' average charge" for vaginal delivery. The latter is presented for 1989 in Exhibit 3 of Schwartz, Colby, and Reisinger (1991), and is deflated or inflated for other years by GKM using data on hospital cost inflation. For 1989, then, this variable can be directly re-calculated, and the values compared to those reported in Table 1c of the original paper. The results are below.

	CA	CO	FL	IA	IL	NJ	PA	WA	WI
Table 1c	0.03	0.12	0	0.33	0.06	0.04	0.11	0.03	0.19
Our Recalculation	0.06	0.25	0	0.36	0.08	0.07	0.17	0.07	0.32