The Effect of Community Uninsurance Rates on Access to Health Care

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Objective. To investigate the effect of local uninsurance rates on access to health care for the uninsured and insured and improve on recent studies by controlling for time-invariant differences across markets.

Data Sources. Individual-level data from the 1996 and 2003 Community Tracking Study, and market-level data from other sources, including the Area Resource File and the Bureau of Primary Healthcare.

Study Design. Market-level fixed effects models estimate the effect of changes in uninsurance rates within markets on access to care, measured by whether individuals report forgoing necessary care. Instrumental variables models are also estimated.

Principal Findings. Increases in the rate of uninsurance are associated with poorer access to necessary care among the uninsured. In contrast with recent evidence, increases in uninsurance had no effect on access to care among the insured. Instrumental variables results are similar, although not statistically significant.

Conclusions. Changes in rates of insurance coverage are likely to affect access to care for both previously and continuously uninsured. In contrast with earlier studies, there is no evidence of spillover effects on the insured, suggesting that such policy changes may have little effect on access for those who are already insured.

Key Words. Uninsurance, access, community effects, spillovers

In a 2003 report, the Institute of Medicine (IOM) hypothesized that communities may be affected by high rates of uninsurance “through the problems that uninsurance poses for local health care institutions and services” (Institute of Medicine 2003), arguing that uninsurance may affect access to or quality of care for both the uninsured and the insured. A 2009 IOM report summarizes recent evidence suggestive of a spillover effect on the insured, stating that “when community-level rates of uninsurance are relatively high, *insured* adults in those communities are more likely to have difficulties obtaining needed health care” (Institute of Medicine 2009). The Patient Protection and Affordable Care Act of 2010 aims to expand coverage substantially, which could have positive effects for those who gain coverage as well as for those who were

previously insured or remain uninsured. If the strain from high rates of uninsurance is concentrated among those providers who disproportionately serve the uninsured, the barriers to care arising from high rates of uninsurance may be more likely to affect the uninsured than the insured.

Existing studies assess that the association between the rate of uninsurance in a market and access to care report mixed findings, with one study finding an effect on access for low- and moderate-income individuals (Brown, Wyn, and Teleki 2000), another finding no effect for low-income individuals when controlling for other market factors (Andersen et al. 2002), and a third finding a negative effect for the insured, but not for the uninsured (Pagán and Pauly 2006). Other recent studies by Pauly and Pagán present evidence of spillovers from community uninsurance rates to the insured for a number of outcomes, including access to care (Pauly and Pagán 2007, 2009). The IOM summarizes these studies as suggesting that uninsurance leads to negative spillovers to the insured, possibly due to lower revenue streams, greater financial strain, and lower demand for quality health care (Institute of Medicine 2009). These results may not reflect the causal effect of uninsurance rates, as the studies rely on cross-sectional data and have not fully accounted for the role of both observable and unobservable market factors in determining access to care.

This article analyzes how variation in uninsurance rates within markets over time affects access to care, specifically whether an individual forgoes necessary medical care, employing models that control for both market and time and using instrumental variables to attempt to address potential endogeneity. This approach improves on previous studies by controlling for time-invariant market characteristics that may confound cross-sectional analyses, and thus may be more relevant for assessing the potential effects of policies that reduce the rate of uninsurance. In contrast to earlier studies, the results suggest that changes in uninsurance rates primarily affect those aspects of the local health care system that influence care for the uninsured.

CONCEPTUAL FRAMEWORK

Whether an individual has difficulty accessing necessary medical care is influenced by both individual-level characteristics (such as gender, race, education,

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income, and insurance) and contextual or market-level characteristics (such as physician supply, hospital capacity, competition, and managed care penetration) (Andersen and Davidson 2001), including the rate of uninsurance in the community. Existing literature indicates that market-level factors affect physician practice patterns and access to care; for example, physician supply is associated with access to care for vulnerable populations and individuals in nonurban areas (Pathman, Ricketts, and Konrad 2006); in markets with higher managed care penetration, low-income individuals have reported more difficulty accessing care, and physicians have been found to provide less charity care (Cunningham 1999; Cunningham et al. 1999; Litaker and Cebul 2003). Measures of access to care, such as whether an individual forgoes or delays receiving necessary medical care, are a function of these individual and market factors. The observable binary outcome will depend on the constraints facing the individual in accessing care. In general, barriers will be lower for those who have insurance relative to the uninsured, and among the uninsured, barriers will be lower for those in communities with more safety net resources per individual.

After controlling for characteristics of the market and the local safety net system, individuals in areas with increasing rates of uninsurance may be less likely to access necessary medical care, although the effect of the uninsurance rate may differ across insurance coverage categories. There are a number of theoretic arguments supporting the hypothesis that individuals experience diminished access in markets where the uninsurance rate is high or increasing. The IOM has hypothesized that the presence of a large uninsured population could lead to lower revenue streams for providers due to lower demand for care or a greater proportion of uncompensated care, and that this could affect both the uninsured and insured (Institute of Medicine 2003). However, in markets with high rates of uninsurance, it is likely that there is greater stress on public hospitals, community health centers, and other safety net providers (Brown, Wyn, and Teleki 2000). Thus, the community effect of uninsurance may not be constant across insurance groups; rather, it may be the case that the effect is greater for the uninsured than for the insured. Furthermore, among the insured, spillover effects resulting from the impact of high rates of uninsurance on the local health care market may be concentrated among those with public insurance such as Medicaid, who may also rely disproportionately on safety net providers.

Pagán and Pauly (2006) hypothesize a number of mechanisms through which the rate of uninsurance may specifically affect access to care for the insured. First, given that providers who serve the uninsured also care for

insured patients, financial difficulties due to uncompensated care may lead them to cut back on the services they offer to all patients. Second, if high uninsurance rates lead to an increase in the price of health care and insured individuals are required to pay proportional coinsurance, this may decrease demand for care among the insured. Third, lower demand for services could shrink the market, leading to reductions in particular types of specialty care. All of these mechanisms may affect access to care for the insured, but they are likely to have a greater effect on access for the uninsured. Providers who serve a relatively large share of uninsured patients might be those who are most likely to cut back on services or close when there is an increase in the rate of uninsurance. While this will affect both uninsured and insured patients, uninsured patients are less likely to be able to substitute other sources of care. In addition, providers may respond by shifting their focus to patients with a reliable payment source and services that are reimbursed generously by insurers. Similarly, potential price increases could disproportionately affect the uninsured who would pay for a greater proportion of care out-of-pocket (as opposed to the smaller fraction of the cost as paid by an insured patient) and may be charged full list prices (Hadley et al. 2008; Melnick and Fonkych 2008). Finally, unprofitable services that may most likely be reduced include those that are used disproportionately by the uninsured, such as emergency department services, AIDS services, psychiatric emergency care, and substance abuse care (Bazzoli et al. 2005; Institute of Medicine 2009). In addition, if there are constraints on capacity, particularly for charity care, they are most likely to affect the uninsured. Each of these mechanisms implies that we might expect a negative effect of uninsurance for all individuals in the market, although we also expect the effect to be greatest among the uninsured.

DATA

Data on access to care come from the Community Tracking Study (CTS), which tracks a cohort of 60 U.S. communities, including 51 metropolitan areas. Data from 1996 to 1997 and 2003, the earliest and latest years available, are pooled (Center for Studying Health System Change 2000, 2006). The main outcome variables of interest are whether an individual reports not having received medical care he or she believed was necessary ("forgoing care") and whether an individual postponed getting necessary care ("delaying care"), conditional on not having reported forgoing necessary care, in the previous year. Measures such as these are typically used to quantify access to care (Berk,

Schur, and Cantor 1995; Cunningham and Kemper 1998; Herring 2005) because they capture realized access to care that the respondent deems necessary.

Relevant covariates in the CTS data include age, gender, education, family income, race, and insurance type (private versus public) as well as self-reported general health. Following Pagán and Pauly (2006), the sample includes adults aged 18–64, because programs and resources for uninsured children are different from those available to adults and were changing more over this period than were programs for uninsured adults (Dubay et al. 2007), and because the response of access for Medicare beneficiaries to changes in the uninsurance rate may differ from the under-65 population. Market-level variables from the CTS, including the proportion of adults who are uninsured (the uninsurance rate) and the percentage of adults who are publicly insured, are calculated for each site-year using appropriate survey weights.

Additional covariates are included in some models to control for changes in the health care safety net over this period. The full set of covariates is not included in the main specification because some of them may be endogenous, particularly over the long run, or may represent a causal pathway through which uninsurance rates affect access. Including them is informative to the extent that it may clarify the pathways through which uninsurance affects access to care. Data on funding for Federally Qualified Health Centers (FQHCs) and public and nonprofit “disproportionate share” hospitals in 1996 and 2003 are used to control for changes in the safety net that might contribute to changes in access to care for the uninsured. The Uniform Data System files from the Bureau of Primary Health Care (BPHC) contain information on FQHCs, including numbers of grantees and some information on revenue (Lo Sasso and Meyer 2006). Number of BPHC grantees in each market and nonpatient revenue from federal and nonfederal grants measure the availability and generosity of FQHCs in the area. Data on Medicaid disproportionate share hospital (DSH) payments to states for various years come from the Federal Register (National Archives and Records Administration 1997, 2005). Medicaid DSH allotments comprise the majority of federal DSH funding (Wynn et al. 2002). As DSH data are at the state level and some CTS sites overlap multiple states, site-level estimates are created by taking a weighted average of DSH spending per capita, weighted by the proportion of the CTS site population in each state based on the 2000 Census (Missouri Census Data Center 2008). All dollar values are adjusted to 2003 dollars using the all-items consumer price index (CPI).

Area Resource File (ARF) data from the American Medical Association Physician Masterfiles on total number of active patient care physicians, as well as active patient care physicians in general practice, general family medicine, general internal medicine, and general pediatrics, are used to construct variables representing the number of primary care physicians (PCPs) and specialists per 10,000 population in an area (Starfield, Shi, and Macinko 2005). These variables and ARF data on the number of hospital beds per capita are used to control for variation in health care capacity across markets over this period.

Finally, data on market characteristics that may be associated with financial pressure on providers in the area are included in some models. Metropolitan statistical area (MSA) level data on managed care from Interstudy are used to construct overall HMO penetration for 1996 and 2002.¹ Managed care penetration is defined as the sum of Medicare, Medicaid, and privately insured HMO enrollees divided by the total MSA population. Hospital competition is measured using the MSA-level Hospital Herfindahl-Hirschman Index (HHI), constructed from American Hospital Association data, based on admissions and adjusted for hospital system ownership.²

Two measures are used to instrument for the rate of uninsurance in the market. First, data on unemployment from the Bureau of Labor Statistics' Local Area Unemployment Statistics File are obtained through the ARF. Second, data on health insurance premiums from the Kaiser Family Foundation/Health Research and Educational Trust survey of employers are used to construct adjusted average premiums measuring health care costs in each MSA. As health insurance premiums are based on expected costs, differences in premiums across areas and changes in premiums over time should reflect differences in the cost of health care in those areas. Premium costs will be affected by the type and generosity of plans offered, as well as firm and industry characteristics. Adjusting for many of these plan and employer characteristics yields a premium measure that more accurately reflects changes in the cost of health care in an area (Chernew, Cutler, and Keenan 2005). Data from 1995 to 1997 and 2002 to 2004 are pooled and matched with the individual-level data from 1996 and 2003, to increase the number of premium observations for a particular time period.

To construct the measure of premium costs, annual premiums for a single individual plan for each policy are regressed on plan and firm characteristics, including plan type (HMO, PPO, POS, or indemnity plan), whether an employer offers multiple plans and the interaction of plan type with the indicator for multiple plans; plan deductible, coinsurance (in network and out of network), and copayment amounts; whether the plan offers prescription

drugs, outpatient mental health benefits, inpatient mental health benefits, and well-baby care benefits; firm size; and industry. Regression coefficients on MSA dummies are used to construct adjusted health care premium costs not explained by plan and firm characteristics.

METHODS

The primary model of interest considers binary measures of access to care as a function of the uninsurance rate in the community, as well as individual characteristics, market area fixed effects, and time-varying market characteristics. I estimate logit models of each of the measures of access to care for individual i , in market area j , at time t , where the uninsurance rate in market j in time period t is the independent variable of interest, and all models control for a vector of individual-level covariates (including time-varying market-level variables in some specifications³), an indicator for year, and market fixed effects. Standard errors are clustered at the site (MSA) level. Models for the uninsured and insured samples are estimated separately. Including market and time, fixed effects controls for any unobservable time-invariant market characteristics that might affect access to care within the market and any overall trend in access. Data from the beginning and end of the period are used to estimate the effect of the net change over the entire 1996–2003 time period, because changes over shorter periods may be subject to more noise, and access outcomes considered may not respond immediately to changes in uninsurance. Changes in the uninsured and insured populations over time may lead to a selection problem if the population changes differentially across markets along dimensions that are not captured by the observable individual-level covariates in the data. To the extent that changes are similar across markets or are captured in the covariates included in the model, the results should not be biased.

As an extension of the main market fixed effects models, I also estimate instrumental variables models with market fixed effects. There may be threats to the exogeneity of the uninsurance rate, including omitted market-level variables that affect both the rate of uninsurance and the availability of care, or reverse causality. Including market fixed effects that account for any time-invariant unobservables and adding covariates that capture individual demographic characteristics and time-varying market characteristics may ameliorate these concerns, although there may be unobservable time-varying factors that affect both the rate of uninsurance and the availability of care in the market. The rate of uninsurance could be a function of access to care, as

well. If access problems in a market reduce the demand for insurance, this would result in poor access leading to higher rates of uninsurance.

I instrument for the uninsurance rate in the market using two of the primary drivers of rising uninsurance as instruments: premium costs and unemployment rates. While one may worry that premiums could be related to access through, for example, changes in cost sharing, after adjusting for plan and employer characteristics (which account for, among other things, changes in cost sharing and benefits that could affect access to care among the privately insured), the adjusted premiums capture the cost of care in the market resulting from local practice patterns and technology use. Furthermore, as the number of excluded instruments exceeds the number of endogenous variables, the adequacy of the instruments can be assessed by testing the over-identifying restrictions. The assumption that health insurance premiums and unemployment rates are correlated with the rate of uninsurance has been established in the literature (Kronick and Gilmer 1999; Cawley and Simon 2005; Chernew, Cutler, and Keenan 2005; Gilmer and Kronick 2005) and is tested explicitly in the first-stage regressions.

Two-stage residual inclusion (2SRI) (Terza, Basu, and Rathouz 2008) is used to estimate IV models with binary outcomes. The first-stage estimates the potentially endogenous uninsurance rate variable as a function of the included instruments and excluded instruments (adjusted premiums and unemployment), including average characteristics of individuals in the market for the relevant sample to account for individual-level characteristics and weighting by the number of observations in the market. The second-stage models include residuals from the first stage. Standard errors are calculated using a bootstrap.

RESULTS

Summary statistics for uninsured and insured individuals in the CTS sample and market-level characteristics are reported in Table 1. Between 1996 and 2003, the aggregate proportion of uninsured reporting having to forgo care is relatively constant, whereas fewer uninsured report delaying care. The proportion of insured either forgoing or delaying care decreases. Both samples are slightly older and include fewer whites and more Hispanics in 2003 relative to 1996, although the changes in race are larger among the uninsured. Changes in income and gender distribution are not significant for the uninsured. When the samples are stratified by markets experiencing an increase

Table 1: Summary Statistics for CTS Insured and Uninsured Samples and Market-Level Factors

<i>A. CTS Variables by Year and Coverage</i>	<i>Uninsured</i>			<i>Insured</i>		
	<i>1996 Mean</i>	<i>2003 Mean</i>	<i>p-Value</i>	<i>1996 Mean</i>	<i>2003 Mean</i>	<i>p-Value</i>
Forgo care	15.9%	15.7%	.801	6.1%	5.7%	.037
Delay care	25.0%	20.7%	.000	19.5%	17.0%	.000
Either forgo or delay	41.0%	36.4%	.000	25.7%	22.7%	.000
Publicly insured	n/a	n/a	n/a	10.9%	13.7%	.000
Age	34.5	36.9	.000	40.2	42.7	.000
Female	50.0%	50.4%	.702	54.3%	56.0%	.000
Years of education	12.2	12.2	.105	13.8	13.9	.000
Family income/poverty level	2.3	2.4	.061	4.2	4.5	.000
White	59.2%	53.3%	.000	77.6%	76.2%	.000
Black	15.9%	16.4%	.543	10.9%	10.9%	.846
Other	4.4%	6.0%	.000	4.3%	5.0%	.000
Hispanic	20.5%	24.3%	.000	7.1%	8.0%	.000
Excellent health	22.8%	19.0%	.000	26.9%	21.2%	.000
Very good health	32.8%	31.9%	.320	39.4%	40.4%	.014
Good health	26.9%	28.6%	.065	23.2%	25.3%	.000
Fair health	13.6%	16.6%	.000	7.9%	9.9%	.000
Poor health	3.9%	3.9%	.890	2.6%	3.2%	.000
N	5,303	4,116		31,702	25,825	
<i>B. Market-Level Factors by Year</i>	<i>All Sites (n = 60)</i>			<i>Urban Sites (MSAs) (n = 51*)</i>		
	<i>1996 Mean</i>	<i>2003 Mean</i>	<i>p-Value</i>	<i>1996 Mean</i>	<i>2003 Mean</i>	<i>p-Value</i>
% Uninsured	14.8%	15.3%	.652	14.7%	15.1%	.755
% Publicly insured	10.0%	11.9%	.017	9.9%	11.8%	.019
Number of BPHC grantees	3.70	4.67	.270	3.92	5.00	.280
Total FQHC grants (nonpatient revenue)	\$10,706,134	\$19,708,592	.038	\$11,975,545	\$22,467,116	.036
FQHC grants per uninsured individual	\$48.50	\$79.91	.084	\$42.86	\$86.33	.026
Total state Medicaid DSH allotment per capita	\$69.85	\$51.34	.023	\$72.43	\$53.85	.043
Primary care physicians per 10,000 pop	6.86	7.59	.019	7.24	7.99	.011

continued

Table 1. *Continued*

<i>B. Market-Level Factors by Year</i>	<i>All Sites (n = 60)</i>			<i>Urban Sites (MSAs) (n = 51*)</i>		
	<i>1996 Mean</i>	<i>2003 Mean</i>	<i>p-Value</i>	<i>1996 Mean</i>	<i>2003 Mean</i>	<i>p-Value</i>
Specialists per 10,000 population	16.19	17.02	.497	18.06	18.91	.413
Unemployment rate	5.3%	6.0%	.033	5.1%	5.8%	.021
Beds per capita	0.0041	0.0035	.019	0.0042	0.0036	.038
Hospital HHI				0.31	0.26	.168
HMO penetration				0.27	0.29	.555
Adjusted single premium				\$3,071	\$4,184	.000

Notes. *p*-values correspond to test of whether means are equal in 1996 and 2003. Sample includes individuals from all 60 CTS sites.

*HMO premium represents 50 site observations; adjusted premium represents 49 site observations.

versus decrease in uninsurance rates, patterns of changes in demographics are the same across the two sets of markets (see Supporting Information Appendix Table A1). Panel B presents market-level means for both 1996 and 2003 for all sites and for MSA sites only. FQHC grants, Medicaid DSH funds, PCPs, and hospital beds, as well as both variables used as instruments, change significantly over this period.

Of the 60 markets in the CTS sample, 31 experienced an increase in uninsurance rates over this period (mean increase of 3.9 percentage points) and 29 experienced a decrease (mean decrease of 3.2 percentage points). Table 2 describes the distribution of uninsurance rates and changes in these rates across the markets in the sample. Among sample markets with uninsurance rates in the lowest quintile in a given year, the mean rate of uninsurance was 8.8 percent in 1996 and 7.0 percent in 2003; the mean rate in the highest quintile was 22.7 percent in 1996 and 25.9 percent in 2003. Among markets with change in uninsurance in the lowest quintile over this period, the mean change was a 5.7 percentage point decrease, whereas markets in the highest quintile experienced a 6.9 percentage point increase on average.

Table 3 reports coefficients and marginal effects from logit regressions, including market and time effects for the uninsured sample. The marginal effects presented in column (2) indicate that a one percentage point increase in the uninsurance rate in the community is associated with a statistically significant 0.44 percentage point increase in the probability of forgoing necessary

Table 2: Variation in Rates of Insurance Coverage across CTS Markets and Changes in Insurance Rates

<i>A. Mean Market-Level Uninsurance Rate by Quintile of Uninsurance Rate</i>				
<i>Percentage Uninsured Quintile</i>	<i>1996</i>		<i>2003</i>	
	<i>Mean</i>	<i>SD</i>	<i>Mean</i>	<i>SD</i>
1	8.8	1.8	7.0	2.4
2	11.6	0.8	10.9	0.8
3	14.2	0.8	14.5	1.8
4	16.9	0.7	18.4	1.2
5	22.7	3.8	25.9	3.6
<i>B. Mean Change in Uninsurance by Quintile of Change</i>				
<i>Change in Uninsurance Quintile</i>	<i>Absolute Change (Percentage Points)</i>		<i>Percent Change</i>	
	<i>Mean</i>	<i>SD</i>	<i>Mean, %</i>	<i>SD, %</i>
1	-5.7	1.9	-40.7	10.2
2	-1.8	1.0	-13.3	5.7
3	0.5	0.8	2.7	4.8
4	2.7	0.9	20.1	5.9
5	6.9	2.6	46.5	18.1

care.⁴ Estimates from the reduced-form model that includes only insurance rates and individual-level covariates, which yields the most conservative results, indicate that moving from an uninsurance rate that is half a standard deviation below to half a standard deviation above the mean is associated with a 2.7 percentage point increase in forgoing care. This result is robust to the inclusion of additional time-varying market-level variables (the estimated marginal effect increases to 0.47), indicating that the effect of changes in these variables over time is accounted for by the market fixed effects. Market-level time-varying covariates (added in columns [3]–[4]) are generally insignificant, and many do not have the expected sign. Columns (5)–(8) present the results of regressions of whether individuals reported delaying care among the uninsured sample who did not report forgoing care. The estimates are consistently positive, indicating that the uninsurance rate is associated with a higher likelihood of delaying care, but are insignificant for all model specifications.

Results from models for both the forgo care and delay care outcomes for the insured sample are presented in Table 4. Coefficients on the uninsurance rate in both the main specifications and the models with

Table 3: Market Fixed Effects Logit Results for Forgoing Care and Delaying Care among the Uninsured Sample

	1 = Reported Forgoing (Delaying) Care/0 = Did Not Report Forgoing (Delaying) Care					
	1	2	3	4	5	6
	Forgo Care			Delay Care		
	Logit	Marginal Effect	Logit	Marginal Effect	Logit	Marginal Effect
% Uninsured	0.0349*** (0.0115)	0.00439*** (0.00145)	0.0379** (0.0154)	0.00474** (0.00192)	0.00527 (0.0115)	0.00847 (0.00331)
% Publicly insured	-1.855 (1.967)	-0.233 (0.247)	-1.142 (2.130)	-0.143 (0.266)	-0.183 (1.498)	-0.575 (2.037)
Age	0.00859*** (0.00277)	0.00108*** (0.000346)	0.00877*** (0.00295)	0.00110*** (0.000366)	0.000345 (0.00215)	0.00037 (0.000417)
Female	0.303*** (0.0564)	0.0381*** (0.00704)	0.310*** (0.0576)	0.0388*** (0.00716)	0.654*** (0.0465)	0.693*** (0.00862)
Black	-0.403*** (0.0851)	-0.0460*** (0.00878)	-0.442*** (0.0872)	-0.0497*** (0.00876)	-0.480*** (0.0937)	-0.482*** (0.101)
Other	0.241* (0.141)	0.0327 (0.0205)	0.145 (0.146)	0.0190 (0.0199)	-0.301** (0.134)	-0.404*** (0.136)
Hispanic	-0.565*** (0.0845)	-0.0636*** (0.00850)	-0.586*** (0.0881)	-0.0659*** (0.00888)	-0.882*** (0.0897)	-0.907*** (0.0913)
Years of education	0.0525*** (0.0155)	0.00661*** (0.00195)	0.0540*** (0.0157)	0.00675*** (0.00195)	0.0647*** (0.0112)	0.0657*** (0.0114)
Family income/poverty level	-0.131*** (0.0165)	-0.0165*** (0.00203)	-0.123*** (0.0171)	-0.0154*** (0.00210)	-0.0383*** (0.0140)	-0.0378*** (0.0154)
Post (year = 2003)	0.0276 (0.0814)	0.00347 (0.0103)	-0.206 (0.258)	-0.0255 (0.0317)	-0.222*** (0.0733)	-0.354 (0.279)
Number of BPHC grantees			0.0490* (0.0285)	0.00612* (0.00356)		0.00349 (0.0261)
FQHC grants per uninsured			0.000133 (0.000792)	1.66e-05 (9.89e-05)		0.000448 (0.000638)

continued

Table 3. Continued

1 = Reported Forgoing (Delaying) Care/0 = Did Not Report Forgoing (Delaying) Care							
Forgo Care				Delay Care			
1	2	3	4	5	6	7	8
Logit	Marginal Effect	Logit	Marginal Effect	Logit	Marginal Effect	Logit	Marginal Effect
Total state Medicaid DSH allotment per capita		0.00190 (0.00305)	0.000238 (0.000381)			-0.000514 (0.00300)	-9.64e-05 (0.000563)
Primary care physicians per 10,000 population		0.194 (0.146)	0.0242 (0.0182)			-0.110 (0.171)	-0.0206 (0.0321)
Specialists per 10,000 population		0.0555 (0.0546)	0.00693 (0.00683)			0.0245 (0.0551)	0.00460 (0.0103)
Beds per capita		89.58 (211.6)	11.19 (26.43)			-171.3 (184.9)	-32.14 (34.68)
Hospital HHI		-0.999 (0.709)	-0.125 (0.0886)			-0.625 (0.713)	-0.117 (0.134)
HMO penetration		-0.136 (0.719)	-0.070 (0.0898)			0.241 (0.643)	0.0452 (0.121)
Constant	-2.477*** (0.306)	-7.108*** (1.562)		-1.513*** (0.276)		-0.626 (1.797)	
Observations	9,419	8,326	8,326	7,913	7,913	7,001	7,001

Notes. Robust standard errors in parentheses (clustered at the site level). All regressions include site effects. Sample includes 60 sites, except regressions including hospital HHI and HMO penetration, which include 51 MSA sites only. For each regression, the first column presents the logit coefficient, and the second column presents the marginal effect of each variable x_j defined as $\partial \Pr(y = 1 \mid x) / \partial x_j$ evaluated at the multivariate point of means, or the marginal effect for an average individual.

* $p < .1$.
*** $p < .05$,
**** $p < .01$,

Table 4: Market Fixed Effects Logit Results for Forgoing Care and Delaying Care among the Insured Sample

	<i>1 = Reported Forgoing (Delaying) Care/0 = Did Not Report Forgoing (Delaying) Care</i>							
	<i>Forgo Care</i>				<i>Delay Care</i>			
	<i>Logit</i>	<i>Marginal Effect</i>	<i>Logit</i>	<i>Marginal Effect</i>	<i>Logit</i>	<i>Marginal Effect</i>	<i>Logit</i>	<i>Marginal Effect</i>
% Uninsured	0.00596 (0.00973)	0.000310 (0.000506)	0.00767 (0.00940)	0.000404 (0.000496)	-0.000536 (0.00593)	-8.30e-05 (0.000916)	-0.00164 (0.000897)	-0.000254 (0.000897)
% Publicly insured	-0.00800 (0.0145)	-0.000416 (0.000755)	0.00320 (0.0181)	0.000169 (0.000954)	-0.00200 (0.00944)	-0.000309 (0.00146)	-0.00265 (0.00983)	-0.000410 (0.00152)
Public insurance coverage	0.616*** (0.0468)	0.0396*** (0.00356)	0.574*** (0.0452)	0.0363*** (0.00346)	0.127** (0.0499)	0.0203** (0.00818)	0.132** (0.0538)	0.0211** (0.00884)
Age	-0.00365** (0.00182)	-0.000190** (9.37e-05)	-0.00392** (0.00196)	-0.000207** (0.000103)	-0.00654*** (0.000888)	-0.00101*** (0.000137)	-0.00704*** (0.000854)	-0.00109*** (0.000132)
Female	0.279*** (0.0369)	0.0143*** (0.00184)	0.279*** (0.0401)	0.0145*** (0.00207)	0.349*** (0.0187)	0.0534*** (0.00281)	0.333*** (0.0183)	0.0510*** (0.00275)
Black	-0.108** (0.0485)	-0.00538** (0.00234)	-0.104** (0.0493)	-0.00527** (0.00242)	0.0322 (0.0452)	0.00502 (0.00710)	0.0542 (0.0468)	0.00849 (0.00743)
Other	0.250*** (0.0887)	0.0144** (0.00560)	0.234** (0.0928)	0.0135** (0.00588)	-0.135** (0.0661)	-0.0200** (0.00946)	-0.171** (0.0668)	-0.0251*** (0.00935)
Hispanic	0.0170 (0.0368)	0.000889 (0.00368)	0.000640 (0.0705)	3.37e-05 (0.0372)	-0.184*** (0.0552)	-0.0271*** (0.00772)	-0.207*** (0.0678)	-0.0302*** (0.00778)
Years of education	0.0262*** (0.00919)	0.00136*** (0.000476)	0.0248** (0.00965)	0.00131*** (0.000507)	0.0247*** (0.00661)	0.00382*** (0.00102)	0.0262*** (0.00704)	0.00404*** (0.00109)
Family income/ poverty level	-0.0943*** (0.00750)	-0.00490*** (0.000374)	-0.0926*** (0.00781)	-0.00488*** (0.000394)	-0.0528*** (0.00491)	-0.00817*** (0.000753)	-0.0502*** (0.00482)	-0.0776*** (0.000740)
Post (year = 2003)	-0.0528 (0.0503)	-0.00274 (0.00260)	-0.0851 (0.101)	-0.00447 (0.00528)	-0.149*** (0.0253)	-0.0230*** (0.00387)	-0.0972 (0.0690)	-0.0150 (0.0106)
PCPs per 10,000 population			0.168 (0.110)	0.00886 (0.00580)			-0.0606 (0.0662)	-0.00937 (0.0102)
Specialists per 10,000 pop			-0.0581 (0.0528)	-0.00306 (0.00278)			-0.0229 (0.0380)	-0.00353 (0.00588)

continued

Table 4. Continued

	1 = Reported Forgoing (Delaying) Care/0 = Did Not Report Forgoing (Delaying) Care							
	1	2	3	4	5	6	7	8
	Forgo Care			Delay Care				
	Logit	Marginal Effect	Logit	Marginal Effect	Logit	Marginal Effect	Logit	Marginal Effect
Beds per capita			105.2 (103.7)	5,539 (5,460)			-34.04 (58.79)	-5,263 (9,091)
Hospital HHI			0.366 (0.391)	0.0193 (0.0206)			0.0161 (0.294)	0.00249 (0.0454)
HMO penetration			-0.353 (0.430)	-0.0186 (0.0227)			0.129 (0.287)	0.0200 (0.0444)
Constant	-2.856*** (0.235)		-3.164** (1.283)		-1.520*** (0.173)		-0.211 (0.939)	
Observations	57,527	57,527	51,248	51,248	54,054	54,054	48,116	48,116

Notes. Robust standard errors in parentheses (clustered at the site level). All regressions include site effects. Sample includes 60 sites, except regressions including hospital HHI and HMO penetration, which include 51 MSA sites only. The delay care models are fit on the subsample of individuals who do not report forgoing care; hence, the sample sizes are smaller than for the forgo care models. For each regression, the first column presents the logit coefficient and the second column presents the marginal effect of each variable x_j , defined as $\partial \Pr(y = 1 | x) / \partial x_j$ evaluated at the multivariate point of means or the marginal effect for an average individual.

* $p < .1$.
** $p < .05$.
*** $p < .01$.

time-varying market covariates for both the forgo care and delay care outcomes are small and statistically insignificant. These results indicate that there is no effect of changes in the uninsurance rate on either outcome for the insured sample. Controlling for self-reported health does not affect the coefficients of interest for either the uninsured or insured samples (see Appendix Table A2).

Additional analyses examining the effect of uninsurance on the probability that uninsured and insured individuals forgo necessary care stratify markets by whether they were above or below the mean rate of uninsurance at baseline and whether they saw an increase or decrease in uninsurance over the period (see Appendix Tables A3–A6). When markets are stratified by baseline uninsurance rate, estimates for the uninsured from both samples suggest a significant increase of approximately 0.3–0.7 percentage points in the probability of forgoing care, indicating that in markets with both high and low baseline rates, increases in uninsurance are associated with greater difficulty of accessing care among the uninsured. Results for the insured remain small and insignificant for both sets of markets. When markets are stratified by whether they experience an increase or decrease in uninsurance, the estimates for the uninsured sample in markets with increasing uninsurance are significantly positive and larger than for the full sample, and estimates for the uninsured in markets with decreasing uninsurance are small and insignificant. Thus, results in the full sample appear to be driven primarily by markets experiencing an increase in uninsurance rates. Results for the insured are again small and insignificant.

Standard diagnostics support the use of the adjusted premium and unemployment measures as instruments for the uninsurance rate.⁵ Results of 2SRI models for the uninsured and insured are available upon request. For the uninsured sample, the logit coefficient on the uninsurance rate in the forgo care IV model with all market-level covariates is similar in magnitude to (slightly larger than) the coefficients from the basic fixed effects regressions, although not statistically significant. Similarly, the coefficient on the uninsurance rate in the delay care model is positive, but insignificant. Coefficients for the insured sample are generally negative, and in all cases considerably smaller than those for the uninsured and statistically insignificant. These results are similar to the fixed effects results in direction and magnitude, although less precise by nature of IV estimation, indicating that as reported in Tables 3 and 4, the effect of the uninsurance rate on access is larger among the uninsured than the insured.

DISCUSSION

This article provides evidence that changes in uninsurance rates primarily affect those aspects of the local health care system that influence access to care for the uninsured. These results contrast with earlier studies, which rely on variation between markets and find that the rate of uninsurance is associated with poorer access for the insured (Pagán and Pauly 2006; Pauly and Pagán 2007, 2009) but does not affect the uninsured (Pagán and Pauly 2006).⁶ Unlike these cross-sectional studies, the methodology employed controls for time-invariant market characteristics that may have confounded earlier analyses. Resulting estimates of the effect of changes in the uninsurance rate within markets on access to care indicate that higher rates of uninsurance are associated with worse access to care for the uninsured, but that there is not a significant effect of the uninsurance rate on access for those with insurance.

There are four main limitations to this study. First, the outcomes measuring access to care are based on self-reported survey data and thus reflect any bias implicit in such data. In addition, the outcome measures do not make any distinction between primary or preventive care, specialty care, and emergency or hospital services. These data are widely used, including in other studies, considering the effect of uninsurance rates. Second, although this research improves on earlier studies by using data from different points in time to examine the effect of changes in the uninsurance rate within markets, the data do not follow the same people over time and thus may be subject to selection bias if the populations of insured or uninsured are changing in ways that are not captured by the observable data and that are associated with access to care. Controlling for self-reported health does not change the results. Third, as in related research, MSAs are used to define markets, although, in some cases, smaller areas may be more relevant geographic units. Finally, the data available from the CTS cover only a 7-year time period, and thus, I am unable to fully assess the longer term dynamics of changing uninsurance rates.

Based on their finding that the uninsurance rate has negative effects on both quality and access for the insured (though not necessarily for the uninsured), Pauly and Pagán argue, “Paradoxically, the insured may gain the most from helping the uninsured if that assistance is not targeted specifically at providing services to the uninsured only” (Pauly and Pagán 2007). As they find evidence of a negative spillover to the insured (e.g., an association between

high uninsurance and poor access to care), they infer that increased support through the expansions in the safety net will reinforce spillovers to the insured population, whereas policies that allow the uninsured to purchase insurance and subsequently use mainstream medical services will improve care for the uninsured without negatively affecting the insured.

In contrast, my findings suggest that changes in the uninsurance rate within a market have no effect on access to care for the insured, but they have a significant effect on access to care for the uninsured. Stratifying markets by whether they experienced an increase or decrease in uninsurance indicates that the results may be driven primarily by increased difficulty in accessing care in those markets experiencing an increase in uninsurance. While estimates in markets with decreasing uninsurance rates are small and insignificant, it may be the case that longer adjustment periods are necessary to observe potential improvements in access associated with decreases in rates of uninsurance. Overall, the results suggest that changes in coverage may affect both those who gain or lose coverage and those who remain uninsured, but they may not have an effect on access to care for the continually insured in the area. Furthermore, if the negative effect on access for the uninsured operates through the reduction in resources per uninsured individual, then increasing safety net resources could also have a positive effect on access for the uninsured, without affecting access for the insured. More research is needed to understand the specific mechanisms through which uninsurance affects local markets and the long-term dynamics of changes in uninsurance.

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NOTES

1. Geographic areas in the Interstudy data changed between 2002 and 2003. Before 2003, data were aggregated to the MSA or primary MSA, whereas in 2003, they used updated core-based MSAs as geographic units of observation. I used the 2002 data for 2003 to maintain consistency in matching CTS sites to appropriate geographic areas.
2. The HMO penetration and hospital HHI measures were used in Scanlon, D. P., S. Swaminathan, W. Lee, and M. Chernew. 2008. "Does Competition Improve Health Care Quality?" *Health Services Research* 43 (6): 1931–51. Michael Chernew shared the aggregate data.
3. Models including the full set of market characteristics do not include individuals from rural sites because some variables were available only at the MSA level.
4. Linear probability models suggest that a 1 percentage point increase in the rate of uninsurance is associated with a 0.47 percentage point increase in the probability of forgoing care. Results for logit, probit, and OLS models are similar for all results presented.
5. F-statistics from first-stage regressions using each of the instruments separately and in combination for the uninsured and insured samples well exceeded the standard threshold of 10. The Sargan statistic is insignificant for all models, lending further support to the validity of the instruments.
6. I replicate Pagan and Pauly using the 2000–2001 CTS household survey and estimating random effects logistic models of forgoing care (unmet need) as they did. The results are qualitatively the same, with the uninsurance rate being significantly associated with unmet need for insured adults, but not for uninsured adults. (Results not reported.)

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SUPPORTING INFORMATION

Additional supporting information may be found in the online version of this article:

Table A1: Summary Statistics for CTS by Year and Coverage, Stratified by Markets Experiencing Increases versus Decreases in Uninsurance Rates.

Table A2: Market Fixed Effects Logit Results Including Controls for Self-Reported Health Status.

Table A3: Market Fixed Effects Logit Results for Forgoing Care among the Uninsured, Stratified by Markets with High Versus Low Baseline Uninsurance Rates.

Table A4: Market Fixed Effects Logit Results for Forgoing Care among the Insured, Stratified by Markets with High versus Low Baseline Uninsurance Rates.

Table A5: Market Fixed Effects Logit Results for Forgoing Care among the Uninsured, Stratified by Markets Experiencing Increases versus Decreases in Uninsurance Rates.

Table A6: Market Fixed Effects Logit Results for Forgoing Care among the Insured, Stratified by Markets Experiencing Increases versus Decreases in Uninsurance Rates.

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