Variations in the Uninsured: State and County Level Analyses

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VARIATIONS IN THE RATE OF UNINSURED, STATE AND COUNTY ANALYSES

VARIATIONS IN THE UNINSURED: STATE AND COUNTY LEVEL ANALYSES

EXECUTIVE SUMMARY

Background

This is a study of why health insurance coverage varies across the country and within a state. While individual characteristics of the uninsured are well known, relatively little research has been devoted to explaining the systematic variation in state and county-level rates of uninsurance. Since these aggregate rates are the targets of many recent policies -- health insurance market reforms, Medicaid expansions, high risk pools -- the policy development process can only be helped by a better understanding of the covariates of rates of determinants coverage and uninsurance.

Policy makers care about rates of health insurance coverage for at least two reasons. One, the uninsured often need and receive expensive acute care to stabilize or cure their medical conditions, and few are able to pay the full costs of these services. These costs are then shifted in various ways onto the population at large. Two, despite these implicit subsidies for expensive care, the uninsured generally use fewer health services than the insured population, and some suffer physical, financial, and emotional consequences from the relative lack of care.

Given the established links between health insurance coverage and access to timely health care, and given the importance of that access to the overall well-being of any population, the sheer amount of variation in rates of health insurance coverage is striking. At the state level, proportions of the overall population that were uninsured in 1995-96 varied from 7.9% in Wisconsin to 24.4% in Texas (Bennefield, 1997). Within the highest overall coverage state of Wisconsin, county level rates varied from 4.0% to 17.0% in 1994. These are the phenomena we are trying to explain.

Specifically, using the state as the unit of analysis, we test for the effects of various state policy interventions on rates of uninsurance, private coverage, and Medicaid coverage for the non-elderly population, while controlling for economic structures, population demographics, and health care service market characteristics. Then, using Wisconsin counties as units of analysis within a common regulatory environment, we test for the relative importance of economic, demographic, and health service market characteristics in explaining overall rates of health insurance coverage.

State Model Results

Using Current Population Survey data from 1989-1995 and building on a data structure originally developed by Zuckerman and Rajan (1997), our model's explanatory power was very good, with over 90% of the variance in uninsurance rates, private coverage rates, and Medicaid coverage rates explained. We found that small group reforms, nongroup market reforms, mandated

benefits, Medicaid policies and certain kinds of high risk pools all affected coverage, though not always in the intended ways.

Specifically for group market reforms, guaranteed issue plus other issue reforms (guaranteed renewal, limits on pre-existing conditions, and portability [credit for prior coverage periods]) were strongly associated with higher rates of private and overall coverage. This is consistent with other ongoing research at the Urban Institute and elsewhere, and comports with the goals of these laws. At the same time, premium rating restrictions in the small group market were just as clearly associated with lower rates of private and overall health insurance coverage, and the two countervailing effects, where present in the same states, effectively cancel each other. Thus states which implemented all four issue reforms plus rating restrictions had no net effect on total coverage in their states. This result is consistent with the inference that rate restrictions, imposed on a pre-reform environment closer to experience rating, lead to more groups observing rate increases than rate decreases. This result is also consistent with what is known about the skewed distribution of health expenditures as well as simulation models of the small group market developed recently for the US Department of Labor by Blumberg and Nichols (1997) and Marquis and Buchanan (1997).

Nongroup market reforms were not adopted by enough states to permit as rich a set of specifications as we explored with group reforms. Nevertheless, our results strongly suggest that guaranteed issue plus nongroup premium rating restrictions in tandem work to decrease overall and private health insurance coverage. Thus, while they surely helped some individuals who are likely to be high risk, state nongroup reforms appear to have decreased coverage.

Other policy findings of note are that mandates for alcohol or drug abuse treatment decrease coverage, that any willing provider and freedom of choice provisions have no measurable net effect on coverage, and that high risk pools decreased Medicaid coverage and may have some effects leading to increased private coverage, though this last result was not completely robust over all statistical tests we ran.

County Model Results

The goal of this part of the study was to test for the relative importance of economic, health market, and demographic factors in explaining overall health insurance coverage within a common regulatory environment, i.e., at the county level in Wisconsin. Wisconsin is an exemplary state to study because it has a low average rate of uninsurance, some very good data on its health service markets, and a unique household survey that permits county-by-county estimates of the rate of uninsurance. We were able to derive statistically satisfactory estimates for 29 Wisconsin counties in 1991 and 1994.

The county models for each year fit less well than did our state model for multiple years, and the unemployment rate was the only significant covariate of county-level uninsurance in both years. As expected, higher unemployment rates were associated with lower levels of insurance coverage.

The size of the unemployment effect was larger in 1994 than in 1991. This may reflect the fact that 1991 was a recession year and in 1994 the economy was strong, so that the marginal job had health insurance attached to it only in the later year.

In 1994, the county model explained 71% of the variance in rates of uninsurance, and health service market variables were also significant in consistent ways. Higher hospital prices were associated with lower rates of coverage (elasticity around -1.0), and a greater supply of physicians increased coverage. This latter effect is consistent with the hypothesis that greater physician supply lowers average physician services prices and premiums though competition.

Surprisingly, a preponderance of employment in small firms did not affect coverage in Wisconsin in either year, and other economic structure variables were either weak, insignificant, or inconsistently significant across years.

County level models are less stable than state level models, but the unemployment variable proves the economy always matters. We tentatively conclude that health service markets can be important in establishing basic ranges of feasible coverage through their impact on prices and premiums, but their marginal influence can be dwarfed by economic factors, especially in a recession year.

Conclusions

States can affect overall rates of coverage though health insurance market policies, but they should be mindful that some policies may counteract others. Specifically, guaranteed issue in the small group market appears to work, but rating restrictions may offset some or all of the gains in coverage. In the nongroup market, none of the policies tried thus far appear promising, if the goal is to increase the rate of overall coverage.

More work needs to be done in refining measures and estimates of the effects of premium rating restrictions. Due to the imprecision of secondary sources, we were unable to calibrate our measures of rating restrictions more precisely than the simple presence or absence of some kind of restriction. For example, looser rate bands, in conjunction with guaranteed issue and like reforms, might be expected to be more conducive to coverage expansions than are tighter rate bands. We were not able to test this hypothesis in this study.

At the county level, unemployment is paramount, but health services markets, especially hospital and physician markets, can significantly affect coverage, at least under favorable economic conditions. More work at this level would be helpful, but should be done only if county sample sizes are large enough to permit many more observations at standard levels of precision than we could marshal in this study.

INTRODUCTION

This is a study of why health insurance coverage varies across the country and within a state. While individual characteristics of the uninsured are well known, relatively little research has been devoted to explaining the systematic variation in state and county-level rates of uninsurance. Since these aggregate rates are the targets of many recent policies -- health insurance market reforms, Medicaid expansions, high risk pools -- the policy development process can only be helped by a better understanding of the covariates of rates of determinants coverage and uninsurance.

Policymakers care about rates of health insurance coverage for at least two reasons. One, the uninsured often need and receive expensive acute care to stabilize or cure their medical conditions, and few are able to pay the full costs of these services. These costs are then partially shifted in various ways onto the population at large, leading to higher prices and premiums. At the same time, the amount of uncompensated care hospitals can afford to provide is decreasing due to competitive pressures on hospitals. Closures and conversions of the nonprofit and public hospitals, which provide most of the uncompensated care delivered in this country, put the uninsured at risk of being denied acute care in the future unless new public funding is forthcoming (Mann et al. 1996; Marsteller, Bovbjerg and Nichols 1998; Weissman 1996).

The second reason policymakers care about the uninsured is that despite these implicit cross-subsidies for expensive care, the uninsured generally use fewer health services than the insured population, and some suffer physical, financial, and emotional consequences from the relative lack of care. We know the uninsured make fewer visits to a doctor, are more likely to use expensive emergency room care, are more likely to be hospitalized for preventable illnesses, and

suffer increased risk of mortality (Blumberg and Liska 1996). In addition to the private gains the uninsured would get if they became insured, society benefits from improving the health of the entire population through productivity improvements and reduced transmission of communicable diseases. In addition, many uninsured people are children, who have no voice in the insurance decisions and who will be healthier throughout their lives if they receive proper medical care when they are young.

Given the established links between health insurance coverage and access to timely health care, and given the importance of that access to the overall well-being of any population, the sheer amount of variation in rates of health insurance coverage is striking. At the state level, proportions of the overall population that were uninsured in 1995-96 varied from 7.9% in Wisconsin to 24.4% in Texas (Bennefield, 1997). Within the highest overall coverage state of Wisconsin, county level rates varied from 4.0% to 17.0% in 1994. These are the phenomena we are trying to explain.

After focusing on the 50 states, Wisconsin is a particularly good state to study in more depth at the country level because the average uninsurance rate is low, it varies so much across counties, and the Wisconsin state government collects many data at the county or local level that are not generally available in other states.

Urban Institute interviews conducted in Wisconsin in April 1997 revealed a variety of theories about how the state achieved such widespread insurance coverage.[1] A majority of interviewees called attention to Wisconsin's very low unemployment rate. In the Madison area,

[1] Interviews were conducted as part of the Urban Institute's Assessing the New Federalism project. See Coughlin et al. (1997).

for example, interviewees estimated that only 1.8 percent of labor force participants were unemployed. With low unemployment, they reasoned, employers must compete for workers on the bases of both wages and benefits. One interviewee felt that the high rate of female workforce participation in Wisconsin, the third highest in the nation, increased the likelihood that any given family will have employer-sponsored insurance. Another response was that unions are relatively strong in Wisconsin,[2] important because unionized workers are more likely to have employer-sponsored coverage (Nichols et al.1997).

Other interviewees suggested that the low uninsurance rate is a result of the corporate culture: employers are "used to providing insurance." The state's employer-sponsored insurance (ESI) rate is high, at 78.6 percent of all people in the state (compared to the national average of 66 percent).[3] This high employer-sponsored coverage rate even reaches the poor and near poor: 22 percent of Wisconsinites in poor families and 57 percent of those in near poor families have ESI, compared to 17.3 and 48 percent nationally.[4]

Finally, some interviewees felt that the spread of managed care had decreased the cost of insurance in the state, leading to wider coverage. Interviewees at the HMO Association reported that the state uninsurance rate had declined steadily as HMO penetration increased, presumably operating through lower insurance prices. In addition, competition among insurers was described as heated, which could be expected to result in more affordable insurance rates.

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[2] In the percentage of workers who are members of a union, Wisconsin ranked 12th in the nation in 1983 and 15th in 1995.

[3] From Urban Institute analyses of 1994-1995 two-year merged Current Population Survey data.

[4] Ibid. Poor families have incomes below 100 percent of the federal poverty guidelines (FPG) and near poor families have incomes below 200 percent of the FPG.

Specifically, using the state as the unit of analysis, we will test for the effects of various state policy interventions on rates of uninsurance, private coverage, and Medicaid coverage for the non-elderly population, while controlling for economic structures, population demographics, and health care service market characteristics. Then, using Wisconsin counties as units of analysis within a common regulatory environment, we test for the relative importance of economic, demographic, and health service market characteristics in explaining overall rates of health insurance coverage.

This report is organized as follows. The next section surveys the literature on determinants of insurance coverage. This is followed by two sections on the state and county models which explain uninsurance as best we can. Finally, we conclude with a brief summary and suggestions for future research.

SECTION I. LITERATURE REVIEW

Why do coverage rates vary so dramatically from one state to another and from one county to another? We review the academic literature on the determinants of health care insurance coverage in five sub-sections: economic conditions (including firm and worker characteristics), demographic characteristics, state insurance regulations, public programs, and market structure and competition.

The Effects of Economic Conditions on Insurance Coverage

Employment is by far the single most important source of health insurance coverage in the US. In 1996 over 64% of the nonelderly (EBRI 1997) and over 61% of all Americans (Bennefield 1997) obtained coverage through an employment-related benefits plan. Therefore economic and employment conditions, e.g., unemployment rates, the percentages of workers in small firms and in certain industries, average wage levels, and worker demographics such as age and marital status are likely to be important determinants of a community's rate of uninsurance.

Despite the obvious connection between employment levels and state and county coverage rates for health insurance, only two empirical studies have actually tried to estimate the relationship in a multivariate context, and their results are not consistent. Diehr et al (1991) used data from nine counties in Washington state to establish that unemployment was a good predictor of uninsurance in 1989. In contrast, Schmidt and Deichert (1996) found no effect of unemployment in 15 Nebraska counties in 1989 and 1991, and they speculated that this could be due to the fact that unemployment was very low in Nebraska in this time period. This is consistent with our general inference from these papers that the determinants of insurance coverage at the county level are likely to vary over time and across states.

While the employment relation is important to health insurance coverage, many people who are employed do NOT have health insurance--indeed, 85 percent of the uninsured in 1996 were employed or dependents of employed persons (EBRI 1997). There are both supply side (firm offer) and demand side (worker preferences) reasons for this, as the research literature has documented. We consider each in turn.

Supply side reasons for uninsured workers. It is well known that small firms are less likely to offer health insurance to their workers than are large firms (Nichols et al. 1997; Jensen and Morrisey 1997, Gabel et al. 1998). Small firms have higher costs than large firms, per worker, of providing health insurance since they have fewer workers over which to spread the fixed costs of administering benefits. This cost disadvantage is exacerbated by higher rates of worker turnover among small firms. Also, with fewer workers, small firms' abilities to spread health risks are impaired, and insurers charge higher premiums because of the extra risk small groups thereby present. The net result of this is that small firms would pay higher premiums for the same benefits, all other things equal. In actual fact, they seem to purchase less generous policies and pay about the same premiums for fewer benefits (Cantor, Long, and Marquis 1995).

Small firms also have shorter life expectancies, experience more variable profits and premiums, and pay lower wages than large firms (Nichols et al. 1997). These characteristics are consistent with less health insurance provision because the uncertainties and opportunity costs of health insurance benefits are greater for small firms.

Demand side reasons for uninsured workers. The fact that small firms offer health insurance less frequently and pay lower wages is a tipoff to the possibility that many workers in small firms may be unwilling to trade their cash wages for health insurance benefits. Small firms

are concentrated in industries and draw upon occupations wherein offering health insurance is not required to attract the workers they need (Nichols et al. 1997). Finally, part-time workers are less likely to be offered and to take coverage from their employers. Since workers implicitly pay for employer-sponsored health insurance through lower wages, some workers are understandably not willing to sacrifice already low money wages for this benefit (Long and Marquis 1993).

While large firms are still much more likely to offer health insurance than are small firms, a recent national household survey found that workers in establishments of *all* sizes were more likely to be offered health insurance in 1996 than they were in 1987 (Cooper and Schone, 1997). Wide notice has been given this study's finding that worker acceptance or take-up rates have dropped by 8 percentage points from 88.3 percent in 1987 to 80.1 percent in 1996. However, since the percentage of married couples in which both the husband and wife work outside the home has increased from 56 to 60 percent since 1987 (BLS 1998), this apparent decline in take-up may be overstated. More families may have two workers who are offered coverage, and therefore the rise in declinations is somewhat artificial. Cooper and Schone report that the percentage of workers with access to employment-based health insurance through one spouse or another held virtually constant over the previous decade, 81.8 percent in 1987 and 82.2 percent in 1996, and that the family take-up rate fell by only 4%, from 93.2 percent to 89.1 percent. Thus, while workers are indeed declining coverage more frequently than they once did, especially at lower wage levels, effective take-up rates have not declined as much as the worker-only numbers have been taken to imply.

The self-employed are somewhat less likely to be insured (75 percent) than other workers (84 percent), and much less likely to have employment-related insurance (46 percent vs. 74

percent, Nichols et al. 1997). This is related to the fact that most self-employed people run businesses with fewer than 10 employees. These rates of coverage imply that proprietors and the self-employed are much more likely to buy nongroup or individual market insurance than are wage and salary workers, probably because they have higher incomes than employees.

The Effects of Demographics on Insurance Coverage

Insurance status also varies by demographic characteristics like age, income, marital status, gender, race, and sometimes geographic region. The most recent national data (from the 1997 CPS) describing the demographic characteristics of the uninsured are shown in Table 1-1. These data make it clear that young adults, those with lower incomes, nonwhites (especially Hispanics), noncitizens, and those who are single are the least likely to be insured. The young may put less relative value on health insurance since they expect to be healthy and their wages are lower than older workers. In multivariate analyses, income is often the most important demographic predictor of lack of coverage (Acs 1995; Comer and Mueller 1992; Frenzen 1993). Of course, age, race, and citizenship are also correlated with income, so that sorting out relative impacts is difficult.

Studies have reached conflicting conclusions regarding the relative likelihoods of coverage of rural vs. urban residents. Most have found that rural residents are more likely to be uninsured (Hartley et al. 1994; Markowitz et al. 1991; Frenzen 1993; and Coward et al. 1993) while one (Comer and Mueller 1992) found that rural residents in Nebraska were no more likely than urban Nebraskans to go without health insurance. Again, this suggests that controlling for state specific influences may be particularly important in generalizable studies of the determinants of health insurance coverage. Markowitz et al. (1991) found in a multivariate analysis that residents of

Western states were the least likely to be insured. Southern states are also likely to have higher than average rates of uninsurance (EBRI 1997).

The Effects of Regulation on Insurance Coverage

There are several types of insurance regulation that might be expected to have effects on coverage through access and price. These include: 1) small group and individual insurance market reforms; 2) selected high-cost mandated benefits; and 3) restrictions on selective contracting between plans and providers. We consider each in turn.

Small group insurance reform & individual reform. Two types of health insurance market reforms may affect insurer offers as well as small business decisions to sponsor and/or individual decisions to purchase coverage: (1) rules of issue (guaranteed renewability, limits on pre-existing condition exclusions, portability or credit for prior coverage, and guaranteed issue); and (2) rating restrictions or controls on premiums charged. These reforms are briefly defined below and described in more detail in Blumberg and Nichols (1995, 1996).

Guaranteed renewability means that insurers cannot refuse to sell to an existing policyholder for reasons of health status or experience. In the absence of other laws, however, they can charge a significantly higher premium. Limits on pre-existing condition exclusions restrict the length of time that coverage can be denied for previously diagnosed problems. Credit for prior coverage, or portability, allows a person who has been continuously insured to count prior coverage periods toward any pre-existing condition waiting period a new insurer may impose. This prevents workers from being forced to go through new waiting periods for prior conditions each time they switch jobs or insurers. Guaranteed issue provisions require an insurer who sells a specific product to anyone to offer to sell that product to everyone else who wants it

(though usually not at the same price).[5] Finally, rating restrictions come in three forms: (1) pure community rating, the same premium for all in a given area; (2) modified community rating, which restricts the factors by which premium quotes can vary (health status and prior experience are not allowed); and (3) rate bands, where the variance or range of premiums allowed to be charged demographically identical enrollees or groups is limited. For example, in Wisconsin, insurers can adjust premiums for age, sex, geographic location, type of business, and benefits design. However, Wisconsin's rate bands limit premium adjustments to no more than 30 percent for experience, health status and duration of coverage. So if the premium for a standard risk 44-year-old male in Milwaukee is $200 per month, the minimum premium for any other 44-year-old male in Milwaukee would be $160 and the maximum would be $240 per month regardless of their specific health statuses. Most states have some form of rate bands.

In general, insurance reforms increase the degree to which insurers are forced to pool health risks across groups and individuals. Less pooling and more risk segmentation is typically more profitable for insurers, but the consequences of unfettered segmentation are potentially volatile premiums for many and some groups and individuals will be unable to purchase insurance from any insurer. These consequences became politically unpalatable in most states in the late 1980s and early 1990s. These market reforms represent a compromise between our collective desire for a free enterprise health system and the wish to structure the small group and individual health insurance markets so that more are better served (at acceptable cost to the currently insured) than in the absence of regulations. States have passed many different combinations of

[5]Prior to 1996, i.e., for the period of our study, most states required only 2 products to be guaranteed issue, rather than all. The federal Health Insurance Portability and Accountability Act of 1996 required that all products in the small group market be guaranteed issue.

these reforms. Many fewer states have passed individual reforms than have passed group market reforms. Table 1-2 reports the reforms which different states passed from 1989 to 1995, our study period.

As Blumberg and Nichols (1996) discuss, scientifically satisfying empirical tests for the effects of insurance market reforms are rare. Generally, multiple reforms are implemented simultaneously, and so causation is difficult to attribute to any specific policy. State-specific baseline premium and benefit package data prior to reform are not likely to be available, rendering comparisons to reformed markets difficult. Transformations of the health care market independent of health insurance reform legislation (e.g., managed care penetration, increased self-insurance and worsening commercial risk pools, the secular rise in the number of uninsured) make it extremely difficult to separate the effects of reforms from general market trends. And finally, most state reforms are relatively new, so most insurance markets have yet to reach a new state of equilibrium, a necessary pre-condition for definitive empirical analysis.

The effects of small group insurance reforms have been studied in two ways. Descriptive analyses typically present simple estimates of the number of insured or of changes in premiums paid in a particular state without controlling for various factors that could explain part of the differences across states and time. These data are derived from available and often different sources at different points in time. The other kind of study is multivariate and multistate, combining coverage or premium observations with person, employer, or state specific "explanatory" variables, including categorical policy variables that at least indicate the presence or absence of specific health insurance reforms. We review the salient examples of each type of study.

Descriptive studies. New York has been studied descriptively, and its experience also serves as a useful cautionary tale. Effective April 1, 1993 and with no phase-in period, insurers were required to use pure community rating in the under-50 group and individual markets. A risk adjustment mechanism (using age, sex, and specific conditions) was implemented.

The first and oft-cited study of these reforms, conducted by the actuarial consulting firm Milliman and Robertson (M&R), estimated that the number of uninsured increased by 405,000 (Litow, 1994). Unfortunately, that study took estimates of the insured before and after reform from two non-comparable data sources -- the Current Population Survey (CPS) and insurer provided data. The CPS, used for the pre-reform estimate, counts individuals who had insurance at some point during the previous year. The insurer data, which were used for post-reform counts, generate estimates of the number of persons with insurance at a specific point in time, a smaller number than those with coverage at any time in the year. In addition, the CPS counts include all employer coverage, whether through commercial insurers or through self-funded plans, while the post-reform counts in the M&R study include only commercial insurers (Institute for Health Policy Solutions, IHPS, 1995). Thus, the M&R estimate of the increase in the number of uninsured caused by insurance reforms is biased upward.

Data collected by the New York Insurance Department show a reduction in the number of those insured through the individual and small group market of 6.8 percent or 88,355 people. Most of these were from the individual market (64,784), with the remaining 23,571 from the small group market (IHPS 1995). Since the small group market is much larger than the market for individual health insurance, we may conclude the bulk of the effect of New York's reforms were felt in the individual market. A considerable number of younger enrollees apparently dropped out

of the New York health insurance market after community rating was implemented. At the same time, coverage fell by considerably less than the M&R report asserted, and perhaps by as much as 75 percent less. About 30 percent of small group enrollees experienced considerable premium increases in one year (20 percent or more), and the number of people experiencing premium decreases was probably smaller than the number experiencing premium increases.

Regulatory entities in Minnesota and California have published studies of their own small group market reforms. In Minnesota, the major provisions implemented in July 1993 included guaranteed issue and renewal, a 12 month limit on preexisting condition exclusions, and restrictions on premium rate variations, as well as other reforms. Since reform, evidence indicates that the number of enrollees in the small group market increased by more than 8-12 percent (Minnesota Department of Commerce 1995). Blue Cross Blue Shield of Minnesota surveys indicate that coverage in the individual market fell by approximately 6 percent (IHPS 1995). Premium increases averaged less than 5 percent in the small group market between 1993 and 1994. These coverage and premium changes are gross figures and the amount directly attributable to the reforms themselves is not known.

The Minnesota experience to date is obviously much more positive than New York's. One important difference is that Minnesota did not impose pure community rating. Another may be the degree of HMO penetration and the resultant state of competition among providers, both significantly more advanced in Minnesota. These conditions, where present, counteract upward pressure on average premiums from rating reforms, perhaps completely.

In California the basic small group insurance reforms, effective as of July 1, 1993, were: guaranteed issue, guaranteed renewal, a 6 month limit on pre-existing condition exclusions, and

premium rating restrictions. The Department of Insurance surveyed indemnity carriers and found that, overall, small group insurance enrollment fell by 28,000 (5 percent) in the first year of reform. The number of small groups insured by indemnity carriers fell by 3,600 (10 percent). Thus, the smallest groups tended to be the ones who dropped indemnity coverage in the aftermath of reform. Most indemnity carrier losses were in urban areas where competition from HMOs is keen (Turem 1995). The California Department of Corporations, which regulates HMOs, found that HMO enrollment in the small group market increased by more than enough to compensate for the decline in indemnity enrollment. (California Department of Corporations 1995). Buchmueller and Jensen (1997) found that the affordability of insurance in California improved between 1993 and 1995 for firms with three to 99 employees. While the median premium did not change, premiums on the high end declined significantly, with no corresponding rise in premiums below the mean. Employer provision of insurance increased significantly among firms with three to nine employees. The authors noted, however, that the effects of small group reforms were indistinguishable from those of a general economic recovery in the state between 1993 and 1995.

New Jersey passed very comprehensive health insurance reforms in 1993 -- guaranteed issue for small groups (2-49) and individuals, standard benefit packages, phase-in of pure community rating (2 years for individuals, 3 years for groups). The reform law also required insurers to either participate in both the individual and the small group markets or to pay an assessment to compensate those who do sell in the individual market. Unfortunately, there is as yet no hard data available to assess the effects of small group reform in New Jersey, though preliminary evidence suggests that the combination of reforms and opportunities did not cause an exodus of carriers and young people, as some had feared (IHPS 1995).

Multivariate studies. Most econometric studies to date have found virtually no effects of insurance reforms. Uccello (1996) found no effect of state reforms and premium taxes on a firm's probability of offering health insurance. Jensen, Morrisey, and Morlock (1995) also found the effects of market reforms to be quite modest. Guaranteed issue legislation was found to have a slightly significant positive effect (at the 10 percent level) on a firm's decision to offer health insurance. Curiously, the presence of tax or employer subsidies for offering coverage had a weakly negative effect (at the 10 percent level) on offering. This result may be the artifact of states with low baseline rates of coverage adopting the subsidy reforms. They also noted that extremely few small businesses knew that the reforms of interest even existed in their states. Additionally, the authors point out, due to the recent implementation of many of the state reforms, it may have been too early to detect effects in 1993.

In a recent summary and extension of their earlier work, Jensen and Morrisey (1997) report that small firms were more likely to offer insurance to their workers in 1995 than in 1993, and that they are much more likely to offer them managed care plans. They suggest that part of the reason may be that small group reforms, especially issue reforms, have served to "level the playing field" for HMOs, in that the reforms reduced the price advantage and selection techniques indemnity insurers have traditionally employed to combat more community-rated HMO products in the marketplace. Guaranteed issue in particular seemed to have a strong effect on the market for firms with fewer than 10 employees. At the same time, enough large firms may have adopted managed care plans and shown that their workers are well cared for that local resistance to managed care by smaller firms may have finally broken down. Jensen and Morrisey also found that the presence of legislation permitting small firms to buy "barebones" benefits packages

increased the likelihood of offer. The authors concluded that small group reforms, taken together, increase the coverage of employees of small firms. Without the reforms, 46.4 percent of firms would have offered insurance, as opposed to the 50.6 percent which actually did offer coverage in the presence of reform. Still, they note that the coverage enhancing effect of small group reforms is likely to be small.

Most recently, Zuckerman and Rajan (1997) used the 1989 to 1995 Current Population Survey (CPS) data to examine the relationship between small group and individual insurance reform and the proportions of state populations that are uninsured. This is the first paper to our knowledge to test for these relationships at the state level. The multivariate regression model used three dependent variables for the nonelderly population -- the uninsurance rate, the percentage of private coverage and the percentage enrolled in Medicaid -- and controlled for state- and time-specific factors with binary variables. One of the authors' most important conclusions is that since specific small group and individual reforms are rarely enacted alone, trying to estimate the separate effects of each of the five types of reforms could yield misleading results because of high correlations among the reform variables. This correlation among the independent variables makes it difficult to identify which ones have truly significant independent effects on the dependent variable.

After analyzing each of the five reforms separately, Zuckerman and Rajan focus on four mutually-exclusive packages of reforms (i.e., those enacted at the same time).[6] Small group packages included implementation of: 1) guaranteed issue, guaranteed renewal, portability, limits

[6] Since the packages are mutually exclusive, there was no problem with correlations among the "package" variables.

on pre-existing condition exclusions and rating restrictions in the same year, 2) all reforms except guaranteed issue, 3) only guaranteed renewal and rating restrictions, 4) and any other set of reforms. Individual reforms were analyzed in the same way, but with only two packages: 1) guaranteed issue, guaranteed renewal, limits on pre-existing condition exclusions and rating restrictions in the same year, or 2) some other collection of reforms.[7]

Using the packages of small group reforms, the models suggest that implementing all five reforms reduces the uninsurance rate by about 0.7 percentage points.[8] This reduction is accompanied by corresponding increases in both private coverage and Medicaid coverage, though the Medicaid effect may be due to other state policy changes implemented at the same time as the insurance reforms.

Individual market reforms, whether analyzed in packages or individually, were found to increase uninsurance rates. The authors conclude that two scenarios are consistent with these findings: either the individual reforms allowed relatively sick individuals to buy policies which raised average premiums for all, or the market rules *per se* caused insurers to exit the individual market, which provided remaining insurers with the market power to raise prices. Given the tiny market shares most commercial indemnity insurers have (Chollet and Kirk 1998), the adverse selection scenario, or at least post-reform premium pricing in anticipation of adverse selection, seems more likely.

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[7] Portability is rarely enacted for the individual market. It is not separated from pre-existing condition exclusion limitations here.

[8] This result was significant only at the 10% level, and thus somewhat less confidence than usual can be attached to it.

In conclusion, evidence on the effect of small group and individual insurance market reforms is still a bit sketchy for hard conclusions, but some patterns are emerging. Guaranteed issue is the only reform which consistently increases coverage and offers of coverage in the small group market. However, guaranteed issue may decrease coverage in the individual market. Comprehensive packages of group reforms have been linked to slightly increased coverage. Finally, an important lesson from the previous literature is that reforms may be better analyzed as "packages" rather than as separable elements since they are passed and implemented in groups.

State mandated benefits. State legislatures have long required that certain health services be covered by health insurance contracts, and sometimes they have required coverage of the services of particular health professionals. Nationwide there are well over a thousand of these kinds of laws, and more are added each year. In recent years legislatures have also mandated specific clinical practices, e.g., specifying minimum length hospital stays pursuant to childbirth. Prior to this recent wave of intervention, states might have mandated that maternity services be covered generally, but would have left it up to the insurers whether to cover particular diagnostic or treatment regimens under that broad umbrella.

The literature has dealt so far with more general benefit and provider mandates. Direct estimates of their effects are difficult, because baseline premium data are so scarce, and because confounding influences on current premiums are so many. Not surprisingly, existing studies have reached different conclusions.

The effect of a specific mandate on *average* premiums, the effect that matters most from a big-picture policy point of view, is the product of a number of variables. These include the fraction of insureds who don't have that benefit now (S_{no}), the marginal effect the new benefit has

on the cost of these kinds of plans (C_{new}), the fraction of insureds who have partial coverage for this benefit now (S_{part}) and the marginal effect on the cost of these kinds of plans (C_{part}). If the mandated benefit affects indemnity plans vs. HMOs vs. PPOs differently, it would be important to distinguish these proportions and differential marginal effects as well.

At a minimum level of complexity, the net premium effect can be expressed as:

$$\text{Net Percentage Premium Increase (NPPI)} = S_{no}*C_{new} + S_{part}*C_{part}$$

So even if C_{new} and C_{part} were large, say 10% and 5%, respectively, if S_{no} and S_{part} were not very large, say 20% and 30% respectively, NPPI could be relatively small, 3.5% in the postulated example. This small but positive average effect even in the face of relatively large premium effects for the S_{no} plans could explain why empirical evidence supporting huge effects, on average, has been relatively weak and, simultaneously, why those who oppose these mandates argue sincerely that the costs to *them* could be substantial.

Two papers, Jensen and Gabel (1992) and Goodman and Musgrave (1988) have found significant negative effects of state mandates on the likelihood of insurance coverage. Jensen and Gabel used data from a 1985 National Federation of Independent Business (NFIB) survey of small firms, and, since the NFIB survey response rate was quite low (17 percent provided data on the survey items relevant to the model) they also checked their model using the spring 1988 Health Insurance Association of America (HIAA) survey of 1,938 firms.

The 1985 model had statistically significant results for two aspects of regulation: the presence of a state continuation of coverage option for terminated workers (which had a negative effect on the likelihood of offering) and the presence of a state mandate for drug abuse treatment

(which had a positive effect).[9] When tested as a group, all of the regulatory variables together had a significant and negative influence on the probability of offering workers insurance at the 10 percent level. The 1988 model also had a few significant individual effects (coverage of psychologists' clinical services, the continuation of coverage option, and the average premium tax rate) and the group of regulatory variables were, again, significantly negative as a group. According to Jensen and Gabel, almost 20 percent of nonoffers among firms in 1985 and over 43 percent of nonoffers in 1988 were attributable to mandates.

Gruber (1994) found two problems with the Jensen and Gabel study which led him to question their results. First, the Jensen and Gabel analysis relies on tabulations of state laws from a Blue Cross and Blue Shield report. Gruber concluded that information in that report was inconsistent with state law in a number of instances. In addition, he notes that Jensen and Gabel did not take into account the fact that after 1986, all firms of more than 20 workers were subject to a continuation of coverage mandate through COBRA. Using Jensen and Gabel's data but with his characterizations of state policy, Gruber finds no effect of state mandates on the decision to offer insurance.

Goodman and Musgrave (1988) used aggregate data to assess the effect of state mandates on the rate of uninsurance among the non-elderly population. They found that mandates were responsible for up to 25 percent of the uninsurance across states. Their analysis is burdened by methodological problems, however. As Gruber (1994) also notes, the number of state mandated benefits may be correlated with any number of a state's policies (Medicaid eligibility, for example)

[9]Although apparently anomalous, a mandate could appear positively associated with the probability of offer if states with higher private sector offer rates are more likely to impose the specific mandate in question.

that have a direct effect on the rate of uninsurance in a state. Not controlling for these other policy variables will likely lead to biased estimates (omitted variables bias). In addition, their measure of the extent of mandates is the total number of mandates, including mandates that insurers merely offer employers certain coverage options. Since the effect of specific mandates on claims costs and premium levels will tend to vary considerably, the mere count of the number of mandates is unlikely to include sufficiently meaningful information for estimation purposes.

In his own analysis, Gruber combines data on workers in firms of 100 or fewer employees from the May 1979, 1983, and 1988 pension and employee benefits supplements to the Current Population Survey (CPS) with data on mandates across states. The focus of his analysis is on state premium taxes and five high cost mandates: alcoholism treatment, drug abuse treatment, mental illness, chiropractic services, and mandated continuation of health insurance benefits for terminated employees and their dependents. His dependent variable is equal to one if the individual has health insurance coverage through their employer and is equal to zero if they do not. The models were run both as linear probability models and probits, the results were not sensitive to the specification used.

Gruber finds no significant negative effects of any of the individual mandates, and the sum of the mandates has a negative but insignificant coefficient. The same is true when the models were modified to control for individual state-specific effects. Gruber also ran models only on those firms of less than 25 employees, those most likely to be sensitive to mandates, and the estimated results were even weaker than those for the under 100 firm size group. In other models, he finds that state waivers of mandates for the smallest firms had no significant effect on

the probability of coverage. This study, carefully done, is powerful evidence against the argument that state mandates have had significant effects on firms' decisions to offer coverage.

Uccello (1996) develops a probit model of a firm's decision to offer and uses data from the 1991 Health Insurance Association of America (HIAA) Employer Survey. She estimated two separate equations -- one for small firms (2 to 49 employees) and one for medium and large firms (50 or more employees). State mandates were measured in two ways: one variable is equal to the total number of mandates in each state, and three specific mandates that have been cited as being particularly expensive, substance abuse coverage, mental health, and psychologist coverage, are included as three dummy variables.

The results of her small firm model show no significant effect of the total number of mandates on the decision to offer. The only significant state mandate measured was psychologist services. By evaluating the regression results at the variable mean, the results indicate that small firms operating in states with a psychologist mandate are 22 percent less likely to offer insurance. The substance abuse and mental health mandate indicators were not significant. None of the state mandate variables were significant in the probit equation for medium and large firms. Unfortunately, since her model was estimated with data from only one year, it is impossible to rule out the possibility that states with psychologist mandates already had the highest levels of mental health insurance coverage and mental health costs.

In addition to their study of small group reforms, discussed above, Jensen, Morrisey, and Morlock (JMM, 1995) included a variable for the number of mandated benefits in the state. They found it to be negative and significant at the 1 percent level, although the magnitude of the effect was apparently quite small. Each additional mandate lowered the probability of offer by six tenths

of a percentage point. However, since the average state has 18 specific mandates (GAO 1996b), this means that on average mandated benefit laws in general were estimated to reduce the probability of small firms offering insurance to workers by almost 11 percentage points.

One way to interpret the differences in findings from the Uccello and JMM studies on the one hand and of the Gruber study on the other is that the cost of mandated benefits may be rising over time. Gruber's study period was entirely within the 1980s. This may be particularly true for mental health benefits, the underlying source of the effect of covering psychologists' services in Uccello's results. In 1995 the BLS reported that while most medium and large firms offered some form of mental health coverage, the percentage of plan participants whose mental health benefits for inpatient care were the same as those for other illnesses fell from 54% in 1980 to 14% in 1993. This must be because employers and insurers perceive the relative cost of mental health coverage to have been rising over this period. Thus mental health mandates in the later part of this time period were probably more expensive than they used to be. This later period is the time frame of the Uccello and JMM studies.

In response to the debate over mental health parity at the federal level in 1996, a number of simulations estimated the potential impact of mental health parity legislation. Two of the most carefully done were by Rogers (1998) of a California law and the Congressional Budget Office (1996) for the federal legislation. Using existing data bases and microsimulation techniques, both studies concluded that premiums would rise, though by a modest 2-4% (much more moderate than many actuarial firms' less careful estimates). Both studies point out that costs would rise less for HMOs since they are more likely to offer parity now than are PPOs and indemnity plans. Both studies also interpret available evidence to suggest that managed mental health care will save

resources compared to unfettered fee-for-service mental health care, and that even PPOs and indemnity plans may turn to managed mental health care in response to parity mandates.

In sum, the evidence to date of the effects of mandated benefits is that they can increase costs, especially various mental health mandates. However, the fact that many firms offer most mandated benefits now, even mental health, means that the effect on average premiums is less than it will be on health plans that offer no or limited versions of the mandated benefit.

Selective contracting restrictions. States may regulate health plan contracts with their enrollees and the contracting arrangements between health plans and providers.[10] Two kinds of laws often enacted between 1989 and 1995 were any willing provider (AWP) laws and freedom of choice (FOC) laws. Proponents of these laws are interested in minimizing what they see as the negative effects of limited networks; detractors believe these restrictions increase the cost of insurance by limiting the ability of managed care plans to contain costs.

AWP laws permit any provider who is willing to accept a plan's terms and conditions to join the plan and serve as a network provider. These laws are intended to maintain patient access to almost all providers, and provider access to all patients, insofar as state law can. AWP laws may apply to an assortment of providers (including hospitals, physicians, pharmacies and nonphysician providers like chiropractors; or all of these) and may regulate a variety of limited-network plan types. The most common AWP laws require all health plans (but often not HMOs)

[10] There is some controversy about whether states can restrict health plan contracts with providers, since states are limited by the Employee Retirement and Income Security Act of 1974 (ERISA) to regulation of the "business of insurance." Some interpret this as permitting regulation of contracts between the insurer and the insured party only. See Marsteller et al. (1997) for additional detail.

to allow any willing pharmacy to join the plan. However, fourteen states' AWP laws protect almost all types of providers and apply to all PPOs, all HMOs or both (Marsteller et al. 1997).

FOC laws allow enrollees to visit any provider they choose, regardless of whether the provider is part of the health plan's network. Some laws allow the out-of-plan use contingent upon the provider's acceptance of the payment rate set by the plan. Others expressly forbid MCOs from assigning financial penalties to enrollees for use of out-of-plan providers.

The central premise behind selective contracting is that managed care organizations (MCOs) can provide high quality care at a lower cost than traditional indemnity insurance plans by limiting the number and balancing the types of providers in the network. MCOs consider the ability to select only certain providers as essential to controlling utilization, lowering costs, and maintaining quality. Thus many observers, including the Federal Trade Commission, believe that AWP laws "...would discourage contracts with providers in which lower prices are offered in exchange for the assurance of higher volume," and "inhibit realization of cost savings, such as reduced transaction and auditing costs, made possible by the ability to contract selectively" (BNA 1994: 2012). Freedom of choice laws could also prove very expensive to MCOs if health plans are not allowed to place any restrictions on enrollees' use of out-of-network care and if the plans cannot differentiate payments for care provided in and out of the network. Since the providers are not required to join the plan under FOC, the MCO has no control over these providers' practice patterns and utilization (Marsteller et al. 1997).

Quantitative analyses of selective contracting restrictions have been limited to simulations of the *potential* costs to health plans (and through them, to consumers) associated with selective contracting restrictions. Generally commissioned by trade associations, these studies all find that

restrictions on managed care contracting increase managed care costs (Wyatt Company 1991; Atkinson and Company 1994; Arthur Andersen 1994, Barents 1998). For example, a Lewin-VHI study concluded that AWP laws, if applied on a national scale, would increase national health spending by as much as $74.7 billion between 1996 and 2002 (Sheils, Stapleton, and Haught 1995). None of these studies, however, offers a comprehensive description of current legislation or differentiates among laws according to their strength. Since not all AWP and FOC laws are the same, quantifying their aggregate impact is speculative. Furthermore, these studies' empirical methods range from nonrandom surveys of actuaries' opinions to multiple regression models with inferences based on coefficients that were not significant at conventional confidence levels. Thus, we consider the current evidence of the effects of selective contracting restrictions on health care costs, and thereby premiums and rates of coverage, to be suggestive, not definitive.

The Effects of Public Programs on Insurance Coverage

Medicaid is the most important public program that covers the health care costs of certain low-income people -- pregnant women, single parents, children, the elderly, the disabled, and the medically needy, whose health care costs make up very large fractions of their incomes. State-sponsored high risk pools are often the only option for another group at high risk of being uninsured, those whose health conditions are sufficiently severe that no private insurer (where permitted) will willingly cover them or their specific condition. These individuals are sometimes called "medically uninsurable."

Medicaid. Several states expanded Medicaid eligibility between 1988 and 1993 in response to federal changes and their own health reform efforts. Thus, Medicaid coverage nationally increased over that time from 8.5 percent to 12.4 percent of the nonelderly population

(Blumberg and Liska 1996). Most analysts have concluded that without Medicaid expansions, higher percentages of low-income people would be uninsured today (Holahan 1997; Cutler and Gruber 1997; Dubay and Kenney 1997). Still, the target efficiency of Medicaid expansions has become a major issue, for most analysts also agree that some individuals who became Medicaid eligible dropped their private coverage to take-up Medicaid because it is free. The difference of research opinion is over the magnitude of this substitution or "crowding-out" of private coverage. Estimates range from 20 - 50 percent, that is from 2 to 5 of each 10 persons who obtained Medicaid coverage during the expansion period are thought to have been previously insured with a private plan and thus do not represent net reductions in the number of uninsured.

The importance of the Medicaid program in insuring low-income populations varies across states. The percentage of population enrolled in Medicaid ranges from a high of 21 percent in Tennessee to a low of 6 percent in Colorado. Each state's Medicaid enrollment level is a function of the state's poverty rate, the generosity of the state's eligibility policies, and the rate at which eligible people actually enroll in Medicaid. Table 1-4 compares the percentages of population that are low-income, Medicaid-eligible and Medicaid-enrolled across states. Comer and Mueller (1992) have argued that one of the reasons for higher uninsurance rates in the southern and western United States is the more restrictive Medicaid eligibility criteria.

State high risk pools. High risk pools are designed to provide insurance to individuals who have trouble obtaining insurance coverage because of some health condition. This group represents only about one percent of people in the United States, but they are expensive (Communicating for Agriculture 1996). High risk pools offer a source of coverage to the medically uninsurable, at high but subsidized premium rates. Risk pool members tend to be the

self-employed, small businessmen, employees of small businesses that don't offer coverage and farmers (Communicating for Agriculture 1996). The size of high risk pools in 1995 ranged from 179 participants in Alaska to 30,470 in Minnesota (Communicating for Agriculture 1996). As a percentage of state uninsured, high risk pool enrollments are typically less than one percent, but may be as low as 0.3 percent or as high as 8.23 percent (Stearns et al. 1997). States with high-risk pools are shown in Table 2-2.

Our review of the literature suggests that no peer-reviewed work has been done on the affects of risk pools on uninsurance rates, though Uccello (1996) found no significant effect of high risk pool assessments on employers' decisions to offer insurance. The relative absence of empirical work is perhaps because most risk pools are so small, and so it seems unlikely that risk pools would have any effect on uninsurance rates. Available studies are concerned with the operation and goals of high risk pools, characteristics of enrollees, enrollment rates and reasons for turnover in the pools, and the escalating costs of financing the pools (Laudicina 1988; Zellner, Haugen and Dowd 1993; Stearns and Mroz 1995; Stearns et al. 1997).

Laudicina (1988) offers commentary on some possible effects of high risk pools, beyond covering a small portion of the uninsured. Officials in Wisconsin and Indiana reported that risk pools were responsible for declines in small businesses' insurance costs because "placing a medically uninsurable employee in the state pool frees [the employer] to obtain standard group health insurance for the rest of their employees" (Laudicina 1988:101). Supporting this notion, Wisconsin surveyed pool enrollees and found that 15 percent of them received contributions to premium costs from their employers (Galanter 1986). Laudicina (1988) reported that officials in Iowa and North Dakota, however, felt that the pools increased the costs of insurance to

employers because of the health premium tax on insurers used to help finance the pool. Some firms subject to the tax (passed on by insurers) will perceive themselves as worse off if they do not have or expect to have any high-risk employees. High risk pools could also decrease costs by lowering uncompensated care burdens for providers. This in turn could reducing cost-shifting to insured patients, private and public alike. So, high risk pools may lower private premiums and Medicaid costs at the same time. Stearns et al. (1997) suggest that there may be an association between the passage of small group and individual insurance reforms and decreases in risk pool enrollment, presumably resulting from increased access to private coverage.

The Effects of Market Structure on Insurance Coverage

The structure of both the health financing and delivery markets would be expected to affect the level of uninsurance in states and smaller areas. As we are using it, market structure includes provider and insurer supply, prices, competition and concentration.

In the insurance market, one would expect that increased entry and competition among health plans could reduce the monopoly power of dominant insurers and thereby lower premiums on average. HMOs have also been found to have a separate, decreasing effect on insurance premiums in some markets due to their cost-cutting practices and ability to negotiate discounts with providers (Feldstein and Wickizer 1995). But competition among health plans could also increase the degree of market segmentation and raise actuarially fair premiums for some groups while lowering them for others. This segmentation effect is most likely if insurance products are very different and appeal to systematically different health risks, that is, if HMOs and indemnity plans are not perceived as close substitutes by consumers. There is considerable evidence that HMOs have enjoyed favorable selection of health risks for some time (Hellinger 1995).

Thus, HMO penetration into markets could either increase or decrease the average premium most people face, especially markets that were formerly dominated by a few indemnity insurers, depending on final market shares. Baker and Corts (1995) present the only known empirical test for the relative importance of the competition effect versus the segmentation effect, using data from a 1991 survey of employers. They find that the market segmentation effect dominates if HMO penetration exceeds 10 to 13 percent, above which higher HMO market share is associated with higher premiums. Their results suggest that greater HMO penetration could reduce net insurance coverage once overall HMO market share reaches a critical mass, at least in the ranges that are observed in their data.

Of course it is possible that HMO-dominated markets would have lower average premiums than would markets dominated by indemnity insurers, for in this case the negative effects of segmentation would be offset by the cost-reducing effects of HMOs. Furthermore, if all plans in a local area were based on similar delivery systems then adverse selection would likely disappear and premium differences would be driven solely by provider payment patterns. As more small firms switch to managed care products, this type of effect may become common (Morrisey and Jensen 1997), but not enough of these kinds of markets existed in the 1991 data used by Baker and Corts for this potential effect to have been observed.

Health service markets comprise the superstructure of insurance markets. Excess bed capacity or high physician supply may result in increasingly competitive service markets, wherein dominant or organized buyers engaged in selective contracting (managed care plans as well as some self-insured employers) could obtain lower health service prices from physicians, hospitals, and other providers. Melnick et al. (1992) found that the largest PPO in California was offered

lower hospital prices in areas with greater hospital competition. If insurance markets are competitive enough to engender lower health service prices for some aggressive buyers, then at least some of these savings are likely to be passed on to aggressive (and large) insurance purchasers through lower premiums. Note that favorable health service prices will not likely be offered to all; small buyers (health plans or self-insured employers) are less likely to benefit.

Given that premiums may move as a result of changes in competition in either insurance markets or the underlying health service markets, how much change in the rate of uninsurance should we expect to observe? This will depend upon the responsiveness of firms and workers to changes in premium prices. Economists measure responsiveness with a concept called "elasticity" of demand, which is defined to be the percentage change in quantity demanded brought about by a 1 percent change in price. Quantity demanded can be measured as the number of insured or the amount of insurance purchased. Typically, higher prices would lead to lower demand for health insurance, and lower prices would lead to increased coverage.

A number of recent studies have estimated the price elasticity of demand on the part of employers, workers, or other individuals. Elasticities may differ among different groups, and the elasticity of demand for purchasing insurance vs. remaining uninsured on the part of individuals will likely be different (and lower) than the elasticity of demand for switching among health plans, given the decision to purchase some plan.

For example, Feldman et al. (1989) and Dowd and Feldman (1994) have found very high elasticities of demand for health insurance among workers. They found that a 1 percent higher premium will lead 4 to 8 percent of workers to switch to lower cost (but similar) plans. But these papers have focused on the demand for switching among specific types of plans. Recently,

Buchmueller and Feldstein (1996) found similar elasticities among workers choosing from a menu of similar plans in a managed competition framework.

But when the plans are less similar, elasticities drop (in absolute value). Morrisey and Jensen (1997) estimated the price elasticity for employers switching between indemnity plans and HMOs to be around -0.33. This is much closer to the elasticities of demand found among firms who were deciding to offer insurance or not, estimated with natural experiment data by Thorpe et al. (1992), also discussed in Helms et al. (1992). Feldman et al. (1997) argue that survey-based evidence may be more representative than subsidy-experiment data, and estimates that the small firm's offer elasticity is between -3.9 and -5.8.

The higher employer elasticities estimated by Feldman et al. are inconsistent with all published estimates of the elasticity of demand by *employees* deciding to purchase or remain uninsured-- including Marquis and Long (1995) and Chernew et al. (1997)--which range from virtually zero to -0.65. One way to reconcile these findings is that employers may be more price sensitive than individual employees when it comes to the binary decision to purchase or not. But given a decision to purchase insurance, recent evidence suggests that both employers and employees are highly price-sensitive when choosing among health plans, at least under conditions approximating managed competition.

To summarize this complex literature, HMO penetration may decrease market prices through spillover effects of its internal cost-reduction methods. In addition, HMOs may reduce the market power of providers and indemnity insurers. But on the other hand, they may increase market segmentation, such that the net effect on average premiums in many marketplaces today is ambiguous *a priori*, though in 1991 they appeared to increase with HMO penetration above 10-

13 percent. It is most likely that some firms will find their premiums decreased while others will pay more for health insurance as HMO penetration increases. There is some evidence that firms may be more price-responsive than individuals when deciding whether to purchase insurance or not, and thus on net we expect the effective elasticity of health insurance coverage with respect to premium changes to be much lower in absolute value than recently estimated elasticities for workers who, given the decision to buy, are switching among plans.

SECTION II. STATE ANALYSIS

States have taken a wide array of approaches to insurance regulation, and state rates of uninsured vary widely. Still it is not clear to what extent these differences are due to state regulatory policies and programs and to what extent they are due to other differences across states, like basic economic conditions, health market structures or demographics.

Model and Hypotheses

The major purpose of our state-level analysis is to examine the effects of state insurance policies and public programs on uninsurance rates. Since these do not vary within states, an examination *across* states is the only way to achieve this objective. The analysis described here builds upon the work of Zuckerman and Rajan (1997), also of the Urban Institute. We have substantially modified their model by analyzing rating restrictions separately from other small group reforms, examining guaranteed issue and rating restrictions in the individual market separately from other combinations of reform measures, adding a range of additional important policy variables, and modifying some of the control variables.

The same variables are regressed in three equations on three dependent variables: percent uninsured, percent with private coverage and percent enrolled in Medicaid. The Medicaid and private coverage equations primarily inform the results of the uninsurance equation and allow us to offer detail on how a variable affecting overall coverage might operate (i.e., through which insurance markets). We expect variables that have negative effects on uninsurance to generally have positive effects in one or both of the other equations. Our model can be summarized as:

$$\{UI; PRIVATE; MEDICAID\} = f(POLICY, PUBLIC, MARKET, ECONOMICS, DEMOGRAPHICS, STATE, YEAR, \epsilon).$$

Where state uninsurance rates (UI), private coverage rates (PRIVATE) and Medicaid enrollment (MEDICAID) are the dependent variables to be explained by vectors of variables representing the presence of certain state insurance regulations (POLICY), public programs (PUBLIC), features of state health services markets (MARKET), economic and employment conditions (ECONOMICS), characteristics of state residents (DEMOGRAPHICS), fixed effects to account for any effects of unmeasured conditions unique to a state or a year (STATE, YEAR), and a residual (ϵ). The following describes the variables represented by each vector, the rationale for inclusion of these variables, and our expectations for their performance in the model.

State Insurance Regulatory Policy. We looked at several state insurance regulations in order to examine their effects on uninsurance.

Small Group and Individual Insurance Reform. The first group of variables concentrates on two different insurance markets--small group and individual. We focused on five major reforms: guaranteed issue, guaranteed renewal, portability, limits on pre-existing condition exclusions and premium rating restrictions. We adopted Zuckerman and Rajan's (1997) general approach of using reform packages and grouped the policies, since states generally pass many of these laws in combination in the same year.

While committed to using reform packages, we wanted to test if premium rating restrictions and reforms governing the issue of insurance could have opposing effects in the small group market.[11] We expect the issue reforms (especially when guaranteed issue is present among them) to improve access to insurance, and hence lower the uninsurance rate, because they make it

[11] We would have liked to do the same for the individual market, but since rating restrictions in the individual market have always been enacted with issue reforms, there were an insufficient number of cases identifying the rating restrictions variable as compared to the issue reforms variables.

harder or impossible for an insurer to refuse to sell to a particular group. Rating restrictions are designed to protect the affordability of insurance for high risk groups and individuals. But what they actually require is that the variance of premiums charged be reduced. If some premiums are decreased, others must be increased if the insurer is not to lose profit. Thus, the net effect of premium rating restrictions in not clear *a priori.* It is possible for the net effect to be either positive on negative on average premiums and thus on overall coverage. Therefore, the net effect of issue plus rating reforms in the small group market is unknowable *a priori* as well.

Fewer states have implemented individual reforms, and their experimentation limits the range of packages that can be meaningfully tested.[12] We compared two types of packages: guaranteed issue with rating restrictions, and all other combinations of individual market reforms (typically limits on pre-existing condition restrictions and guaranteed renewal). It should be noted that the number of states and years where individual reforms are present is small for each of the variables and thus, we are hesitant to make definitive statements about these results. However, state reform packages which feature guaranteed issue and rating restrictions appear to increase uninsurance (.01 level).[13] So do packages consisting of any other combination of issue and rating reforms, taken as a group (.05 level), although the significance of this result falls below conventional levels when California is removed.[14] These results imply that individual market

[12] When variables were defined in a manner similar to the small group reforms, there was an insufficient number of independent cases to identify the rating restrictions variable adequately. Specifically, the variable for the four major issue reforms was different from the rating restrictions variable in only 14 cases. Rating restrictions were different from the other issue variable in 16 cases.

[13] This result is identified by only 4.6 percent of total observations, n=16.

[14] This result is identified by only 4.2 percent of total observations, n=15.

reforms, as they have been enacted in states thus far, have not been successful in expanding access to insurance; in fact they have tended to decrease overall coverage.

Restrictions on Managed Care Plans' Ability to Limit Provider Networks. As indicated in the literature review, several questionable empirical studies along with compelling economic logic suggest that freedom of choice and any willing provider laws which reduce managed care plans' ability to select and control providers are likely to increase costs and hence the price of insurance. If this occurs, then uninsurance rates will increase. No previous studies, to our knowledge, have tested for the effects of selective contracting restrictions directly on uninsurance rates in a multivariate, fixed effects context. We will do so even as we note that since most AWP and FOC laws are limited in the total number of insureds and providers they affect (Marsteller et al. 1997), however, they may not have much of an effect on insurance status.

Benefits Mandates. Many insurers and employers argue that premiums are increased because of costly mandated benefits. If so, they would also decrease coverage. We will test for the effects of the four most expensive mandated benefits: treatment for alcoholism, treatment for drug abuse, general mental health, and chiropractic care.

Public Programs. Other insurance-related state policy decisions include the eligibility thresholds for public programs, and whether or not to offer the programs at all. Such programs offer free or less- expensive coverage options for lower-income state residents.

Medicaid. The Medicaid program is the largest state-administered public insurance program, and despite federal guidelines and minimums, eligibility criteria vary from state to state. More generous Medicaid eligibility will increase Medicaid enrollment and could reduce uninsurance, since Medicaid covers low-income people who often lack alternative sources of

health insurance. However, as eligibility is broadened to higher and higher income levels, alternative sources are sometimes available, and some of the Medicaid expansion could come at the expense of private coverage.

High Risk Pools for the Medically Uninsurable. High-risk pools are designed to help people obtain coverage whose medical conditions make them unattractive risks to insurers. In the absence of state sponsored and subsidized high risk pools many individuals would go without coverage either because no insurer will accept them or because they cannot afford the high premiums private insurers would charge. Without a high risk insurance pool, some portion of these individuals may qualify for the Medicaid program as they spend down their personal resources to pay for their health care needs. Risk pools do tend to be very small, however, as enrollment is often capped, thus it is not clear whether they are large enough to have much of an effect on a state's overall uninsurance rate. However, their very existence may enable insurers to worry less about adverse selection and thereby price lower in the in vidual market which could increase private coverage.

Control Variables

The next three subsections describe the variables that are included in the model primarily as control variables. As such, we offer only very brief hypotheses for these variables' effects in the equations, because these are not the focus of this study.

Market Structure and Competition. These may be the variables, among all the controls, that are most likely to change over time within a state. States with consistent regulatory policy could experience changing uninsurance rates as competition, supply and prices, in both provider and insurer markets, change.

HMO Penetration. HMOs can lower costs and health service prices, but they may also increase average premiums if the risk selection effect discovered by Baker and Corts (1997) is strong enough.

Physician Supply. We expect higher physician supply to be associated with higher insurance coverage. Physician supply is an important control because physicians may locate in an area where there is a high insured rate because of the higher demand for care associated with coverage. A growing supply of physicians could also lead to a decrease in average premiums if service prices fall because of competition among physicians.

Hospital Bed Supply. Hospital beds are another measure of the supply of health services within a state. If a state has an excess capacity of beds, insurers can negotiate lower prices with hospitals, potentially reducing premiums and uninsurance rates.

Hospital Price. Hospital prices may serve as a proxy for the cost of health care within a state. We expect that if provider prices are high, the cost of insurance would also be high, leading to lower overall coverage. Likewise, if provider prices are low, then insurance prices would be lower, and coverage should be higher.

Employment and Economic Conditions. Employment and economic conditions, including firm and worker characteristics and the unemployment rate, have important influences on uninsurance in a wide range of studies.

Unemployment. We expect lower levels of unemployment to be associated with higher rates of health insurance coverage.

Full-time Workers. Since full-time workers are more likely to be offered insurance and to have higher wages, states with larger percentages of full-time workers are expected to have lower than average rates of uninsurance.

Importance of Retail and Services Industries. Workers in the service and retail industries are less likely to be covered by health insurance, at least partly because of worker characteristics. People who work in these industries are often part-time, are paid low wages and tend to change jobs frequently. The jobs may also require lower skills and may attract a less-educated group of people. As more workers are employed in the retail and services industries, the uninsurance rate is likely to climb.

Importance of small firms. Regardless of industry, small firms are less likely to offer health insurance than are large firms. They also tend to pay lower wages. We expect that a larger percentage of workers in small firms will tend to increase the uninsurance rate.

Demographics. Certain characteristics of state residents may also influence state uninsurance rates, so it is important to include demographics as control variables in any model of uninsurance. If we did not include them, we might inaccurately estimate the effects of other variables. Thus, consistent with the literature, we account for: income levels, the age of the population, education levels, marriage rates, families with young children, the size of the nonwhite and Hispanic populations, and the size of the rural population for each state each year.

Data

Sources. The data sources used in this analysis are national surveys and secondary legislative reviews.

Current Population Surveys. The primary data sources for the state-level uninsurance analysis are the March Current Population Surveys (CPS) for 1990 through 1996 (data years 1989-1995). The CPS interviews on average 57,000 households a year, yielding data for between 130,000 and 158,000 civilian individuals not living in institutions. The sample size has declined over time. CPS provides data on insurance status, age, sex, race, education, work status and income. It is designed to be representative of both states and the nation.

We used a variety of secondary sources to construct policy variables describing small group and individual insurance reform (Ladenheim 1995; GAO 1995; IHPS 1995; Laudicina et al. 1996), any willing provider (AWP) and freedom of choice (FOC) laws (Marsteller et al. 1997), state mandated benefits (Laudicina et al. 1996; Gruber 1994), and high risk pools (Communicating for Agriculture 1996). Sources were checked against each other wherever possible. Some information was verified through calls to either the publishing organization or to relevant state officials.

Other data sources for the state-level analysis include: the Bureau of Labor Statistics for median income and unemployment rates, InterStudy for HMO penetration rates, and the American Medical Association (AMA) and the American Hospital Association (AHA), respectively, for physician supply and hospital supply and expenditures.

Measurement Issues. Individuals over 65 years of age were removed from the database because Medicare provides this age group with nearly-universal insurance coverage. In addition, all people in families with an active member of the military were also removed (approximately 2000 unweighted people). Comparison of Medicaid enrollment between the CPS and Health Care Financing Administration (HCFA) records shows that the CPS under-reports the number of

people covered by Medicaid, so the Urban Institute corrects for this problem using a microsimulation called the Transfer Income Model (TRIM2). There were also some changes in the wording of the insurance status questions in the 1995 and 1996 Current Population Surveys; however, the fixed time effects technique should account for these differences.

Variables. Construction of most variables was straightforward. The policy variables are all in the form of 0-1 dummy variables which measure, for each year, the presence or absence of each type of regulation in each state. In addition, high risk pools are also represented by dummy variables, which indicate the operation in the state, in each year, of such a program. Medicaid eligibility policy is measured as the percentage of population eligible in each state according to the Urban Institute's TRIM microsimulation model (Giannarelli 1992). All policy variables were coded to represent implementation dates, where possible, or a year after enactment, where implementation dates were unknown. Control variables are continuous and represent percentages of workers or population, respectively, for the economic and demographic measures.

Descriptive facts and statistics. Tables 1-2 and 1-3 showed each state and the years that various state regulatory policies were in effect during our time period. While some states, such as Michigan and Hawaii, only have mandated benefits, other states, including Louisiana, New York and South Carolina, have a range of regulatory policies. Every state has at least one policy in effect.

Table 1-2 shows that by 1995, 35 states had passed guaranteed issue, guaranteed renewal, portability and limits on pre-existing condition exclusions for the small group market and 45 states had some form of rating restrictions for that market. Most of these policies came into effect in the early 1990s. The individual market issue reforms also began to take effect mostly in 1993 and

1994, however, these reforms are much less common than in the small group insurance market. Only 8 states had passed the four major issue reforms in the individual market by 1995 and 11 states had individual rating restrictions in effect.

Table 1-3 shows other state regulatory programs as well as the public programs. Twenty-four states had implemented a medium or strong version of an any willing provider or freedom of choice law by 1995.[15] Many states passed any willing provider laws in the 1980s. States typically implemented freedom of choice laws in the early 1990s, and any willing provider laws resurfaced again in the 1994-1995 time period.

As for mandated benefits, some were implemented as early as the 1960s, but the table shows only the years they were in effect during our time period. They are widely enforced across the states, especially the mandated benefit for chiropractic care. With the exception of Idaho and Wyoming, every state had at least one of the four most expensive mandates in place by 1995. In 1995, 42 states mandated a chiropractic care benefit and 28 states had a drug or alcohol mandate in place. Only 17 states in 1995 had a mandated benefit for mental health. Missouri and Virginia repealed some of their mandated benefits in the early 1990s.

High risk pools are not as typical among the states as some of the other regulatory policies are. Twenty states had an uncapped high risk pool in place by 1995, while 4 states had a risk pool that limited enrollment. Some of these programs are still fairly new.

Table 2-1 shows the mean and standard deviation for all the continuous variables in the model, including the dependent variables. Means for 0-1 binary variables are shown as the

[15]Marsteller et al. discuss the characterization of states' selective contracting laws. In general, weak laws apply only to pharmacies or represent minor impediments to managed care plans' freedom to contract.

percentage of observations with a value of one. It is important to keep in mind that the mean is taken over a period of seven years, from 1989 to 1995.

Tables 2-2 and 2-3 show state-by-state values for 1989 and 1995 (the beginning and end of our time series) for selected variables. Table 2-2 describes the three dependent variables. The majority of the population is covered by private insurance. Changes in private coverage are consistent with declining overall rates of private coverage in the United States. Vermont and New Mexico saw the largest declines in private coverage between 1989 and 1995. Medicaid coverage nationally increased over the study period by about 3.5 percentage points, reflecting Medicaid expansions over this time period. The largest increase occurred in Tennessee, whose Medicaid enrollment grew from 10.1 percent of the population in 1989 to 21.5 percent in 1995.
While national uninsurance rates increased slightly over this period, some states saw increases and others saw decreases (for Michigan and Connecticut, uninsurance remained exactly the same). Alaska, Tennessee and New Hampshire saw uninsurance decline the most; Ohio, Maryland and Iowa saw the largest increases over this period.

Table 2-3 shows data for the HMO share of the insured population, hospital expenses per adjusted patient day and the percent nonwhite and Hispanic, each for 1989 and 1995. Reflecting the changes in the health insurance market, most states' HMO penetration increased a great deal over the seven years. Many states' penetration rates doubled. For example, Oregon's penetration rate rose from 30 percent in 1989 to 60 percent in 1995.

Each state, over the seven years, also experienced an increase in hospital expenses per adjusted patient day. Washington state had the largest absolute increase, in dollar value, between 1989 and 1995, although Rhode Island's expenses increased by the same percentage, 77 percent.

The last two columns, for percent nonwhite and Hispanic, reflect the percentage of people who are Asian, Native American, African-American, and all other nonwhite persons, as well as all persons who reported Hispanic origin, regardless of race. Some states did not see much change in the size of this population over time. Others had a lower percentage in 1995 than they did in 1989, such as Indiana, Colorado and Louisiana. West Virginia's nonwhite and Hispanic population dropped by half between 1989 and 1995. Still others like Hawaii, California, Alabama and Nevada saw increases in their nonwhite and Hispanic populations.

Empirical Methods

We specified a cross-sectional time series analysis using seven years of data, from 1989 to 1995. The same independent variables are regressed against three dependent variables. Each dependent variable represents an aggregation to the state level of individual-level data from the CPS. The percentages uninsured, privately covered, and Medicaid enrolled are all measured as proportions, bounded by one and zero. Thus, weighted least squares of the natural log of the dependent proportion variable divided by one minus the proportion, or group logit, is an appropriate estimation technique that produces efficient and unbiased estimators (Greene 1997). Weighting each observation by a function of population and the proportion accounts for the latent heteroscedasticity in the grouped data.[16]

In addition, the fixed effects specification includes dummy variables representing all but one state and all but one year of the time series in order to control for any effects that may be unmeasured by our included independent variables. Fixed effects estimation dramatically improves

[16] This procedure produces minimum chi-squared estimates of the parameters that have been shown to have the same properties as maximum likelihood estimates (Amemiya 1985).

the reliability of model results by removing variation that, in the absence of fixed effects, might be incorrectly assigned to variables in the model.

For 1989 through 1995, we estimated (with time subscript suppressed for simplicity):

$$\text{Log}(UI/1\text{-}UI) = \alpha + \beta_1 \text{POLICY} + \beta_2 \text{PUBLIC} + \beta_3 \text{MARKET} + \beta_4 \text{ECONOMICS} + \beta_5 \text{DEMOGRAPHICS} + \beta_6 \text{STATE} + \beta_7 \text{YEAR} + \epsilon$$

where

UI =	Percent uninsured weighted by state population
POLICY =	Presence of small group issue reforms: guaranteed issue, guaranteed renewal, portability and limits on pre-existing condition exclusions;
	Presence of all above small group issue reforms except guaranteed issue;
	Presence of any other combination of small group issue reforms;
	Presence of any small group rating restriction;
	Presence of individual guaranteed issue and rating restrictions;
	Presence of any other combination of issue or rating reforms;[17]
	Presence of any individual rating restriction;
	Presence of strong or medium AWP or FOC laws;[18]
	Presence of a benefits mandate for alcoholism or drug treatment;[19]
	Presence of a benefits mandate for mental health treatment;
	Presence of a benefits mandate for chiropractic care;
PUBLIC =	Percent eligible for Medicaid;
	Presence of a high risk pool with no enrollment cap;
	Presence of a high risk pool with an enrollment cap;
MARKET =	Hospital expenses per adjusted patient day;
	Hospital beds per 100,000 people;
	Physicians per 100,000 people;
	HMO penetration among the insured;
ECONOMICS =	Unemployment rate;
	Percent workers employed in firms with fewer than 25 workers;
	Percent workers employed in retail and services sectors;
	Percent workers employed full-time;
DEMOGRAPHICS =	Median income in thousands;
	Percent persons aged 46-64;

[17] We defined only two mutually exclusive issue reform packages in the individual market rather than three because of insufficient observations to reliably disaggregate policy packages further.

[18] If a state has a weak version of these laws, for example, a prohibition against contracting with mail order pharmacies exclusively, then the state was given a value of zero.

[19] The alcoholism and drug treatment mandates were combined because of high collinearity between the two when measured separately.

	Percent college graduates;
	Percent married;
	Percent of families with a child < 6;
	Percent nonmetropolitan population;
	Percent nonwhite and Hispanic population;[20]
STATE =	49 dummy variables, coded as 1 for each state and zero for all others;
YEAR =	Six dummy variables, coded as 1 for each year and zero for all others;
α =	A constant;
β_n =	Coefficients of the variables; and
ϵ =	An error term, normally distributed with mean zero.

The same model was also regressed on the percentage of people with private coverage (PRIVATE) and on the percentage of people with Medicaid coverage (MEDICAID).

Regression Results. Table 2-4 shows results of the three regression equations. The fit of the model is extremely good: the variables explain roughly 91, 95 and 94 percent of the variance in the uninsurance, private coverage and Medicaid equations, respectively. The state and year fixed effects account for only 11 percent of the variation in the uninsurance model and 7 and 5 percent of the variance in the private coverage and Medicaid equations, respectively.[21] This indicates that the remaining 80, 88 and 89 percent is explained by our policy, program, and control variables.

Nevertheless, we regard some of our policy results as less robust than others, either because very few states implemented the policy, or because they changed when a single state was dropped from the analysis. California was used to test the sensitivity of our results because it is an unusual state with high HMO penetration and high uninsurance rates. It is also one of the

[20] This measure includes Asians, Native Americans, African-Americans, and all other nonwhite persons, as well as all persons who reported Hispanic origin, regardless of race.

[21] These percentages represent the comparison between the explained variation of the model with and without the state and year dummy variables.

relatively few states that enacted both small group and individual insurance market reforms. This section relates the findings of the three regressions in order.

Uninsurance. In the uninsurance equation, the policy and program variables that have significant effects are the small group and individual insurance reforms, mandated benefits for drug and alcohol treatment and Medicaid eligibility. For small groups, the presence of all four of the major issue reforms (guaranteed issue, guaranteed renewal, portability and pre-existing condition exclusion limitations) has a strongly significant, *negative* effect on uninsurance (.01 level), or a positive effect on coverage. When guaranteed issue is not present, but the other three issue reforms are, uninsurance is also reduced with marginal significance (.10 level).[22] Other combinations of the issue reforms (mostly states with guaranteed renewal only) are also marginally significant (.10 level) and negatively associated with uninsurance rates. These results focus attention on guaranteed issue, within the context of the packages that are typically passed, as the major reason for the quantitative impact of the "all four issue reforms" variable (coefficient -0.1107) being so much larger than the other issue packages (-0.0673 or -0.0768).

On the other hand, the presence of any small group rating restriction significantly *increases* the uninsurance rate (.01 level). This result is consistent with the interpretation that rating restrictions, by making premiums more equal across groups, increase premiums for more groups than they decrease premiums for. Furthermore, for the groups whose costs increase, the increase appears to be large enough to induce some of them to drop insurance coverage. Finally, this evidence implies that a larger number of people in groups decline or lose coverage than gain

[22] This result is based on approximately 5 percent of the observations, n=17.

insurance or find alternative coverage sources wherever small group rating restrictions are present.

Thus, in states with both small group issue reforms and rating restrictions, the different types of reforms appear to have countervailing influences on the uninsurance rate. Comparing the coefficients on these variables allows us to measure the relative strength of these countervailing effects. If enacted with all four issue reforms, the coverage-reducing effect of rating restrictions is virtually the same size as the coverage-enhancing effect of the issue reforms. We cannot reject a test that the sum of the coefficients is equal to zero. Thus we conclude that states which enacted guaranteed issue, renewal, portability, limits on pre-existing condition exclusions, and premium rating restrictions will see no change in uninsurance over time.

The same is true when rating restrictions are enacted with other combinations of issue reforms (the sums of the coefficients are again not significantly different from zero). Results for the small group insurance market reforms are highly significant despite a fairly high correlation between the four issue reforms variable and the rating restrictions variable (see Table 2-5).[23]

The individual insurance reforms were tested using a different configuration from the small group insurance reforms.[24] We compared two packages: guaranteed issue with rating restrictions, and all other combinations of issue and rating reforms. It should be noted that the number of states and years where individual reforms are present is small for each of the variables and thus,

[23] This collinearity is caused by the tendency of states to enact rating restrictions in combination with issue reforms.

[24] When variables were defined in a manner similar to the small group reforms, there was an insufficient number of independent cases to identify the rating restrictions variable adequately. Specifically, the variable for the four major issue reforms was different from the rating restrictions variable in only 14 cases. Rating restrictions were different from the other issue variable in 16 cases.

we are hesitant to make definitive statements about these results. However, state reform packages which feature guaranteed issue and rating restrictions are associated with decreased coverage or increases in the rate of uninsurance (.01 level).[25] So are packages consisting of any other combination of issue and rating reforms, taken as a group (.05 level), although the significance of this result falls below conventional levels when California is removed.[26] These results imply that individual market reforms, as they have been enacted in states thus far, have not been successful in expanding overall access to insurance; in fact they have tended to reduce net coverage.

Benefits mandates for drug and alcohol treatment appear to have a weakly significant effect that reduces overall coverage (.10 level). While most firms may not drop coverage because such a mandate is enacted, the mandate may contribute to generally higher premiums over time. These higher premiums lead some people to drop their employer-sponsored or individual coverage.

Medicaid eligibility decreases uninsurance levels (.05 level). This finding is not surprising since the purpose of the Medicaid program is to offer insurance to certain categories of low-income people who might otherwise go uninsured. The marginal effect of Medicaid eligibility on insurance coverage is not particularly large, however. A one percent increase in people eligible for Medicaid yields a twelve one-hundredths of a percentage point decrease in the rate of uninsurance.

In addition to the policy and program variables, several of the control variables have strongly significant effects. Among the non-policy variables that decrease coverage are the

[25] This result is based on only 4.6 percent of total observations, n=16.

[26] This result is based on only 4.2 percent of total observations, n=15.

percent of population that is nonmetropolitan and the percent of families with children under age six (.01 level). The finding that rural populations are less likely to have insurance is consistent with much of the literature. The percentage of people in families with children under may serve as a proxy measure for some other effect. For example, if families with children under six tend to have young parents working in lower-wage jobs, they may be less likely to be insured than families with older children and hence, older parents with higher-wage jobs.

Despite its consistency with the findings of Baker and Corts (1997), many may find it surprising that larger HMO market shares are associated with lower overall coverage, for this indicates that throughout this time period the risk segmentation effect on average market premiums outweighed the cost reducing effects of HMOs. Perhaps because this result is the net of two conflicting forces, the marginal effect of HMO share on uninsurance is small. A one percent increase in HMO market share among the insured yields a one tenth of one percentage point *increase* in uninsurance. Nevertheless, the result that HMO penetration increases uninsurance proved robust to a range of alternative specifications.[27] In addition, it showed a significant effect despite being correlated with a number of other variables (see Table 2-5). Such collinearity normally tends to lower apparent significance.

Other control variables that are associated with lower uninsurance (or higher coverage) include median income (at the .05 level), percent married and percent nonwhite and Hispanic

[27] Specifically, we attempted to identify omitted variables that could bias the coefficient (and its sign) on the HMO share. In this vein, we added percent nonmetropolitan population and state subsidized insurance programs to the model; however, neither of these additions has any effect on the HMO share result. In addition, we removed a series of states from the analysis, seeking outliers that were perhaps driving the result. The result held true despite the removal of as many as five states with high uninsurance and high HMO penetration (including California) and five states with low uninsurance and low HMO penetration.

(both at the .01 level). These first two effects are consistent with expectations, but the last is not. This effect appears to operate through increased Medicaid enrollment (the Medicaid equation is described below).

One interesting point about the rest of the variables in the model is that all the measures of economic conditions except for median income, i.e., firm and worker characteristics, and health service market structure were insignificant. This is true even of unemployment, which is very important in the county model, as discussed below. The state and year fixed effects appear to largely account for the differences these variables might otherwise detect, since many of these variables do not vary much over time within a state.

Private Coverage. The results of this equation are interesting for two primary reasons. First, they offer insight into how variables may affect uninsurance (i.e., *via* changes in private insurance or *via* changes in public insurance). Second, it may provide useful evidence for states that wish to lower uninsurance rates but prefer not to do so through public program expansions.

Not surprisingly, many of the significant variables in the uninsurance model are also significant in the private coverage model and as expected, the relationships run in the opposite direction. Specifically, the results for the small group and individual market reform variables are consistent with the results of the uninsurance model, although the magnitudes of the effects are smaller. In the small group market, guaranteed issue, renewal, portability and limits on pre-existing condition exclusions significantly increases private coverage and rating restrictions decrease private coverage. Unlike the uninsurance model, however, the other packages of issue reforms did not significantly influence private coverage rates.

Individual market reforms also support the findings of the uninsurance model, by significantly reducing private coverage. However, the private coverage equation is more susceptible than the uninsurance equation was to changes in the results when California is removed. Without California, the significance of the individual guaranteed issue and rating restriction variable declines slightly. The significance of the other combinations of reforms variable drops substantially, giving us little confidence in that particular result.[28]

Three of the other policy and program variables are significantly associated with private coverage. First, mandated benefits for alcoholism and drug treatment have significant negative effects on private coverage (.05 level), but other mandated benefits were not significant. The marginal effect of this variable is large: where alcohol or drug treatment mandates exist, private coverage is two percentage points lower. This effect was weakly significant in the uninsurance model. Thus it appears that at least some of the people who lose or drop private coverage because of this benefit do not obtain alternative forms of coverage in its place.

Medicaid eligibility significantly decreases private coverage (.01 level). A one percent change in the population eligible for Medicaid reduces private coverage by 0.2 percentage points. Thus these results are consistent with the findings of moderate levels of crowd-out of private coverage for Medicaid enrollment. Of course, some people may lose coverage involuntarily, and those who lose coverage involuntarily may have become completely uninsured, in the absence of Medicaid expansions. The marginal effect of Medicaid eligibility is larger in the private coverage equation than it is in the uninsurance equation.

[28] The t-statistic for the other combinations of individual reforms variable drops from -2.91 to -1.363.

The presence of a high risk pool without an enrollment cap also has positive effects on private coverage (.05 level). The creation of an unlimited pool for the medically uninsurable appears to *increase* private coverage rates by one percentage point. This is consistent with the inference that insurers are sufficiently reassured against the threat of adverse selection by the existence of an uncapped high risk pool that they price their insurance products lower than they otherwise would have. We cannot emphasize this result, however, because its significance declines dramatically when California is removed from the model.[29] This clearly merits further multivariate research.

Any willing provider and freedom of choice laws do not have significant effects on private coverage, which suggests that they do not have substantial effects on the overall price of insurance. This is perhaps not surprising, since they regulate only a portion of the insurance market, namely the fully-insured business of managed care organizations. Furthermore, they often regulate only some types of managed care plans, with respect to varying provider types. Suspecting that these laws might have stronger effects in states with high HMO penetration, we interacted this dummy variable with the HMO share of the total population. The new term was also insignificant, however. These findings do not necessarily mean that restrictions on plans' ability to contract selectively with providers have no effect on managed care organization prices. But any such price effect is sufficiently small that no net decreases in coverage occur, at least during the time period studied.

Non-policy variables with significant effects on private coverage rates mirror those in the uninsurance model: median income and percent married increase private coverage, while families

[29] The t-statistic drops from 2.097 to 0.812.

with children under six and percent nonmetropolitan decrease private coverage. Unlike the uninsurance results, the percent nonwhite and Hispanic was not significant. This supports the inference that these variables work through Medicaid coverage. Also, the percent of people in the 46-64 age group decreased private coverage (.10 level). Like the policy variables, these variables also influence private coverage in the opposite way they influence uninsurance. In the private coverage model the marginal effects of these control variables tend to be roughly equal to the size of their effects in the uninsurance model. However, the percent married has a larger effect in the private coverage model.

Medicaid Enrollment. Just as the private coverage model can help add detail to the results of the uninsurance model, so can an analysis of the effects of the same policy, economic, market and demographic variables in relation to Medicaid enrollment. Furthermore, states may benefit from assessing the effects of their other policies and programs on Medicaid in order to better coordinate them.

The most interesting policy result of this equation is that the presence of a high risk pool without an enrollment cap has a highly significant, decreasing effect on Medicaid enrollment (.01 level).[30] Where uncapped high risk pools are created, Medicaid enrollment is 1.3 percentage points lower over time, suggesting that access to a high risk pool may help the medically uninsurable avoid spending down to Medicaid eligibility. This results implies an important role for high risk pools in helping limit Medicaid enrollment that states may not have foreseen.

None of the small group or individual insurance market reforms have significant effects on Medicaid enrollment, which suggests that people affected by private reforms do not generally

[30] This finding is robust to the removal of California.

have Medicaid eligibility. Mandated benefits, any willing provider laws and freedom of choice laws are also insignificant to Medicaid coverage.

Medicaid eligibility has a strongly significant positive effect on Medicaid enrollment, as one would expect (and hope). The effect is not quantitatively large, however: a one percent rise in the percentage of people eligible for Medicaid increases Medicaid enrollment by a little more than 0.2 percentage points. A state's Medicaid eligibility policy is only one of several determinants of enrollment, of course. It also depends on state efforts to inform eligibles about their options and the rate at which eligible people decide to enroll. In addition, as mentioned above, as eligibility standards permit people with higher income levels to enroll in Medicaid, take-up rates among the expansion populations decline.

Among the control variables, the percent married decreases Medicaid enrollment (.01 level), although its marginal effect is small at one tenth of a percentage point. The model suggests that Medicaid enrollment is slightly higher over time where there are larger percentages of nonwhite and Hispanic Americans (.01 level). A ten percent increase in the nonwhite and Hispanic population over time leads to a 1.4 percentage point rise in Medicaid enrollment.

In conclusion, it is evident that state small group and individual insurance reform have significant effects on uninsurance rates across states. In the small group context, the effects of rating restrictions appear to run counter to the effects of issue reforms, with issue reforms increasing coverage and premium rating restrictions decreasing coverage. All of the combinations of individual reforms that have been enacted by states so far appear to increase uninsurance. Other insurance policies, like AWP and FOC laws and mandated benefits as enacted are not strong influences on uninsurance rates.

The models also showed two other important results: broader Medicaid eligibility policies decrease uninsurance levels, and high-risk insurance pools decrease Medicaid enrollment and may increase private coverage as well in some states. The effect of high risk pools on Medicaid appears not to translate directly into changes in uninsurance, however.

Limitations. Although we tried to avoid using policy variables with small numbers of observations, in some cases we had no choice. Any results that are based on a small percentage of data points should be regarded with some level of skepticism. Of chief concern are the two individual insurance issue reform variables, which occur in 15 and 16 cases, and the small group reform variable for guaranteed renewability, portability and pre-existing condition exclusions without guaranteed issue, which only occurs 17 times (out of 350 observations). We should not over-emphasize the results associated with these small numbers of values. Of course, state policy, not data collection or analytic strategies, determines how many observations of particular policies there are.

In addition, one potential problem with using CPS data for this kind of analysis is the small sample sizes in some less-populated states (for example, Vermont had less than 1,000 people in two years of the sample). We tested our results to see if they were affected by the small samples in some states by dropping all states with fewer than an average of 1,700 observations per year. This represented the average annual sample size for over half the states. Our results were largely unaffected. As mentioned above, the weighted least squares specification prevents small samples from driving results by weighting the dependent variable by state population.

Also, changes in the CPS questionnaire, sampling design and data collection methods have somewhat limited comparability of findings across years. Most significant among these changes is

a new set of insurance questions inaugurated in 1995, which differ from the earlier versions of the CPS, mostly in the way that individual coverage and employer-sponsored insurance are described (Swartz 1997). We mitigate the possible effects of this difference somewhat by using private coverage as a whole. In addition, the year fixed-effects should control for changes in the CPS data (among other unmeasurable, time-related effects).

Finally, some level of difficulty and error is inherent in summarizing state policies that are enacted with different strength and detail across the states. The impact of these variables may suffer from the necessarily broad brush used to describe complex policies with binary variables.

SECTION III. COUNTY ANALYSIS

Insurance coverage rates vary across regions and counties within a state, just as they do across states. Since regulatory policy is constant at the county level, in this analysis we concentrate on other factors that may explain rates of uninsurance, such as economic and employment conditions, health service market structure, prices and competition. Importantly, a county-level analysis allows us to use measures of health service market competition -- among hospitals, physicians, and HMOs -- that are not available at the state level. This allows us to test for additional determinants of health insurance coverage at the local level.

Model and Hypotheses

This analysis is intended to identify the covariates of uninsurance for counties within the state of Wisconsin. Our model can be summarized generally as:

$$UI = f(MARKET, ECONOMICS, DEMOGRAPHICS, \epsilon).$$

Thus, in our model the county-level uninsurance rate (UI) is explained by vectors of variables representing features of county health services markets (MARKET), employment, firm and worker characteristics (ECONOMICS), characteristics of county residents (DEMOGRAPHICS), and a residual (ϵ). Demographics are included primarily as control variables; the variables of primary interest in this model are the MARKET and ECONOMICS vectors. The following describes in general terms the variables represented by each vector, the rationale for inclusion of these variables, and our expectations for their performance in the model.

Health Care Market Structure Variables. This vector describes the health services markets in the counties, including provider supply and prices and measures of competition among insurers.

Physician Supply. The number of physicians in a county provides a general measure of the supply of health care in the area.[31] We postulate that a higher number of doctors per capita in an area facilitates competition, thus reducing the cost of service, lowering insurance premiums and the uninsured rate. It may also be the case that doctors tend to locate in areas where demand for care (health insurance) is high. Either way we expect that on average counties with a higher number of doctors relative to the population will tend to have lower uninsured rates.

Provider Prices. Hospital and physician prices measure dimensions of the cost of health care in a county. If physician prices are higher in some counties than in others, the level of insurance prices may also be higher, since expenditures on physician services comprise about 20% of national health spending (Levit et al. 1996). This in turn could influence levels of uninsurance. The same is true for hospital prices. We expect insurance premiums to move with hospital costs because hospital expenditures make up almost 40% of total health care expenditures (Levit et al. 1996)). If insurance premiums increase with hospital costs, then health insurance will become less affordable and the uninsurance rate will likely increase.

HMO Penetration. As we discussed in the literature review and in the state analysis section, HMOs can lower costs and health service prices by managing care better or eliciting provider discounts, but they may also increase average premiums if the risk selection effect discovered by Baker and Corts (1997) is strong enough. Whether the selection effect is likely to

[31] Hospital bed supply is a typical measure of the availability of hospital care. Excess capacity among hospitals would tend to lower the price of health care, and potentially premiums and uninsurance rates. Because hospital beds are highly correlated with physician supply in our Wisconsin data, their effects are indistinguishable in empirical work, but including both variables will bias statistical tests of each variable's significance. Thus, we decided to omit hospital beds per capita.

be stronger at the county or the state level is an empirical question on which we have no *a priori* hypothesis.

HMO Concentration. In conjunction with high penetration, if enrollees in a county are disproportionately enrolled in one or two HMOs, this concentration could create market power. This market power could negate any cost-saving effects that managed care techniques might achieve. We expect HMO monopoly power then would lead to increases in uninsurance rates.[32]

Economic and Employment Variables. These include measures of unemployment and firm and worker characteristics across counties.

Unemployment. Because of the importance of employment related insurance, we expect that a low unemployment rate would also mean a low uninsured rate. In addition, the unemployment rate is a reasonable proxy for the general economic health of a community, and we expect that more widespread economic well being would be associated with a higher insurance coverage rates.

Importance of Small Firms. Numerous studies using national data have shown that employees of smaller firms are less likely to be offered insurance and have higher rates of uninsurance generally. We hypothesize that the higher the percentage of workers in small firms in a county, other things equal, the higher the uninsured rate will be.

Importance of Retail and Services Industries. The retail and service sectors of the economy are noteworthy for their low offer rates to their employees (Nichols et al. 1997). In addition, these jobs tend not to pay as well as other employment sectors, making it less likely that

[32] This of course would be true of any dominant insurer, whether HMO or not. However, measures of market share for other types of insurers are not generally available.

retail and service employees will take coverage that might be offered, nor is it likely most could purchase health insurance on their own. Thus we expect that counties with higher proportions of jobs in these sectors will on average have higher rates of uninsurance.

Proprietors. On average, proprietors tend to have higher incomes than those earning wages and salaries (Nichols et al. 1997). As such we expect that proprietors would have higher health insurance coverage rates than those earning wages and salaries because in general, higher incomes are correlated with higher rates of health insurance. Since proprietors have higher rates of insurance coverage, if they are counted as workers in the small firm variable they could inflate the coverage rate among those in small firms (ninety percent of the self-employed own firms with between zero and nine total workers). Thus, accounting for proprietors serves to control separately for such an effect.

People Working on Farms. Agricultural establishments have the lowest health insurance offer rates of any major industry, with only 36 percent offering insurance in 1993 (Nichols et al. 1997). Farms are obviously concentrated in rural areas, however, substantial variation in the percentage of farm workers exists across rural areas. By including a variable for metropolitan areas in conjunction with this farm variable we will test for differences between agricultural rural economies and non-agricultural rural economies.

Demographic Variables. This vector includes important characteristics of county residents that are known to affect insurance status.

Income. Data from the Current Population Survey (CPS) show an uninsured rate of 7.6 percent in 1996 among those in households with incomes greater than $75,000 whereas those living in household with incomes less $25,000 had an uninsured rate of 24.3 percent in 1996

(Bennefield 1997). Thus, we expect that as the median income in a county rises the uninsured rate will decrease.

Rural versus Urban Areas. Previous research suggests that differences in health insurance status exist between urban and rural areas even after controlling for other demographic and economic effects. Since we will control for age distribution, median income, and physicians per capita in our multivariate work, any remaining influence of rural on coverage may reflect differences in taste for health insurance or the perceived value of health insurance.

Size of the Nonwhite and Hispanic Populations. Numerous studies measuring the patterns of uninsurance have identified race and ethnicity as key explanatory variables. Based on this evidence, nonwhites are expected to have lower rates of private coverage but higher rates of Medicaid enrollment, so the net effect on overall coverage is ambiguous.

Data

Sources. Data for the county-level analysis were collected from a wide array of state, federal and private sources. The State of Wisconsin has a strong data collection infrastructure in place. Sources include the annual Wisconsin Family Health Survey (WFHS), a comprehensive telephone survey yielding data for nearly 7,000 individuals, the Wisconsin Hospital Fiscal and Annual Hospital Surveys, which are similar to the American Hospital Association (AHA) Surveys, and the biannual Wisconsin Physicians Profile Survey. We also obtained HMO data from the Wisconsin Insurance Commissioner's yearly publication entitled *Managed Health Care Plans in Wisconsin*. In addition, Blue Cross and Blue Shield United of Wisconsin (BCBSUW) provided us with data on physician prices.

Federal sources include the Bureau of the Census for population and demographic data and the County Business Patterns (CBP), the Bureau of Labor Statistics (BLS) for unemployment, and the Regional Economic Information System (REIS), administered by the Bureau of Economic Analysis, for other labor market variables.

Measurement Issues. As with the state database, we removed individuals over 65 years of age from the WFHS, since seniors almost all qualify for Medicare and are thus virtually all insured. In addition, some observations had missing values for the important health insurance coverage variables. Instead of dropping these observations, we imputed health coverage status for those missing values from similar individuals in the sample.[33]

While the WFHS data sets provided a binary variable for the respondent's insured status for 1992 through 1995, we had to construct this variable ourselves for 1990 and 1991. After consulting with state staff on the structure of the eight-question hierarchy used to determine insurance status, we devised an algorithm to replicate the variable. This procedure may have introduced additional statistical "noise" into the 1991 model.

Another measurement issue is that the Wisconsin Family Health Survey was designed to produce regional and state estimates, not county estimates. After sorting individuals into the counties where they live, we identified two barriers to our analysis. First, while the samples accurately reflect the total household population in various regions of the state, in any single county, they had too many or (more often) too few people to represent the true county

[33] Imputation was based on a random assignment of the insurance status of other individuals in the sample of the same age, sex, race and poverty status. We excluded any observation which had missing values for both health insurance coverage and any one of the other variables. Excluded observations totaled less than 100 over the six years of survey data.

population. Second, in most counties, the sample sizes for any given year of the WFHS were not large enough to make reliable conclusions.

We attempted to correct for the representation problem by reweighting the observations to represent the true population in each county. We post-stratified the sample by sex and three age groups (19 and under, 20-34, and 35-64) to correctly represent the distribution in each county.[34] The new post-stratification weights are the ratio of the number of each type of person in the inflated sample to the population estimate for that type of person in 1991 and 1994, according to the Bureau of the Census.

Post-stratification weighting introduces additional statistical "noise." This means that the weighting scheme may have increased standard errors and lowered the significance of our estimates. This is preferred, however, to reporting estimates based on data that we know to be unrepresentative. Aggregated uninsured rates calculated using our post-stratified weights differed by 1 percent in the 1991 period and 0.8 percent in the 1994 period from the aggregate uninsured rates calculated with the original WFHS weights.[35] We felt these differences were a small enough price to pay for having data more representative of county level rates of uninsurance.

The second problem, insufficient sample sizes, required that we increase the sample for each county. This was accomplished by merging three years of data to represent one period in time. Thus, the 1991 period uninsurance rates represent average values of Wisconsin Family

[34] These weights were applied after the survey's original weighting scheme had accounted for non-response, the sampling rate in each region, households with more than one phone, the time of year that the respondent was surveyed, and finally, a factor to inflate the sample's number of respondents to the total regional and state household populations.

[35] The uninsurance rates compared were averages for the 29 states included in the analysis.

Health Survey data for 1990, 1991, and 1992. Similarly, the 1993, 1994, and 1995 data were concatenated to form 1994 averages. Other works have also used this three-year merging technique, for insurance estimates and for poverty rates (Winterbottom et al. 1995; Haveman et al. 1991; DOC 1991). We limited the mergers to three years since insurance status is less likely to vary dramatically over shorter merged periods. It also allowed us to preserve two distinct time periods (given six years of available data).

Even after merging three years of data, some counties still did not have enough observations to make reliable estimates. Table 3-1 details the mean standard errors and mean widths of the confidence intervals for varying sample sizes. The table shows that as the sample sizes of the counties increase, the mean standard errors and mean widths of the confidence intervals decrease at a rapid pace.

After some experimentation with sample sizes, we set 150 observations as the cut-off for including a county in the regression analysis. We felt that this threshold best balanced estimation accuracy with necessary degrees of freedom. A total of 29 counties had at least 150 observations in *both* periods.

Variables. For the most part, the construction of the variables is straightforward. There are four independent variables, however, where the construction was more complicated than simply taking a percentage. These include the hospital price, physician price, HMO Index and workers in firms with fewer than 20 workers. These are detailed below.

Physician Price Index. To measure the cost to insurance companies of physician services, we created a relative physician price index using 1991 and 1994 data provided by Blue Cross & Blue Shield United of Wisconsin. We used total volume and payments after discounts for 27

common procedures administered to BCBSUW enrollees, in all plan types, across the 29 counties (see Table 3-2). The index, centered around one, ranks each county's average overall price relative to the 29-county average. As relative prices rise, the index has a higher value. We expect that as the index increases, insurance premiums will also grow and result in a higher uninsured rate.

Casemix-Adjusted Hospital Price. We used the casemix-adjusted inpatient net revenue per discharge for general medical and surgical (GMS) hospitals to measure another dimension of the cost of health care in a county. First we had to convert fiscal year data into calendar year data. We adjusted the inpatient net revenue per discharge to account for casemix differences across hospitals by comparing the average length of stay in each inpatient service area (ISA) of a hospital to the statewide average for each ISA.

An important note about this variable is that we had to impute a fairly large number of values for hospital prices in 1991. Computing this price requires having both discharges and inpatient days for each ISA at each hospital, but the 1991 Wisconsin Annual Survey of Hospitals was missing data for at least one of the two necessary ISA data points for about 25 percent of GMS hospitals. For most hospitals, we imputed the missing data points for *either* the number of discharges *or* the number of inpatient days for an ISA based on statewide averages. For example, if a hospital had 500 discharges in the intensive care unit but the number of inpatient days was missing we computed the number of inpatient days by multiplying 500 times the statewide average length of stay for intensive care units. A handful of hospitals had missing data for both discharges *and* inpatient days in a particular ISA. To impute these data we used the distribution of discharges and inpatient days from 1994 for the same hospital to compute the missing data. While

these imputations allowed us to proceed without the loss of data, the method likely imparts additional statistical noise to the 1991 model.

HMO Index. The available state data provide the number of HMO and point of service (POS) plan enrollees (POS data were only availably for 1994) by county and by carrier. We measured HMO penetration as the percentage of county residents enrolled in HMOs divided by the total insured population from the WFHS. HMO concentration was measured as the percentage of the HMO market held by each HMO in a county.[36]

To economize on the number of variables in the model we combined measures of penetration and market share to create an HMO "index." We constructed this index by multiplying HMO penetration by the Hirschman-Herfindahl Index (HHI) for HMOs in each county. The HHI measures market power by summing the squared share of the HMO market held by each HMO operating in the county. Where HMO penetration and concentration are both high, we expect that HMOs would drive up premium levels, thus increasing the uninsured rate. However, lower levels of the index are difficult to interpret.[37]

To address this problem, we turned the index into a dummy variable where a value of one represents a value of the index over 900, a natural break in the data, and a value of zero represents other values. We expect that values of one will tend to be associated with higher levels

[36] We would have rather had market share of the entire insurance market, but these data were not available.

[37] They may be due to low HMO penetration with high concentration (which would likely have no effect on uninsurance) or conversely, high penetration with low concentration (which may lower premiums and hence, uninsurance). We expect different outcomes for these two cases, diminishing the value of using the index as a continuous variable.

of uninsurance, for there a few HMOs have both market power and large overall insurance market shares.

Percentage of Employees in Firms with less than 20 workers. This variable measures the importance of small firms in a county since small firms are less likely to offer insurance than large ones. A number of surveys report a significant increase in offer rates at the level of 25 employees. The closest intervals in the CBP county-level data were at 20 employees and 50 employees. We chose 20 workers or less as our threshold and computed the percentage of the total workforce in firms of this size.

Due to confidentiality concerns, at the county level, the County Business Patterns reports only the number of establishments in each size category, not the number of workers. Therefore we calculated the average number of employees per establishment in each size category at the state level, and assumed that these averages persisted throughout the counties. Thus, if there are 10 establishments with one to five workers in Green County, and the state average number of workers in establishments of size 1 to 5 is 1.6, then Green County is reported as having 16 workers in firms of size 1 to 5.

Descriptive facts and statistics. Table 3-3 presents the mean, standard deviation, minimum and maximum values for each of the variables used in the regressions. The table shows that the uninsured rate decreased modestly from an average of 10.6 percent across the 29 counties in the 1990 to 1992 period to a mean of 9.9 percent in the 1993 to 1995 period. The uninsured rate is only an estimate, however, and is subject to sampling error. Table 3-4 shows the uninsured rate estimate for each of the 29 counties in both time periods along with the associated 95 percent confidence interval for each estimate. As one would expect, on average, the smaller the sample

size in a county, the larger the confidence interval for that county. Some of the confidence intervals are quite large, for example estimates for Marinette County have 95 percent confidence intervals bounded by 8.2 and 18.4 in 1991 and 11.0 and 23.0 in 1994. Table 3-4 also displays the actual sample sizes for each county in each period.[38]

Table 3-3 also shows a number of interesting changes in the independent variables. For example, the mean unemployment rate in 1991 was 5.4 percent and dropped to an average of 4.6 percent in 1994. This is not too surprising because the country was in the midst of a recession in 1991. However, it appears that the recession was particularly damaging to some counties while leaving others largely unaffected. For example, as Table 3-5 shows, Rock County had an unemployment rate of 12.5 percent in 1991, which fell to a rate of 5.2 percent in 1994.[39]

The casemix-adjusted hospital inpatient price per discharge also changed substantially from 1991 to 1994, increasing from near $3,700 to about $5,000. These means, however, mask the fact that dramatic changes in these hospital prices were experienced only in some counties. For example, anecdotal evidence suggests that between 1991 and 1994 the health care market in Walworth County evolved from a characteristically rural market into an extension of the high-priced Chicago health care market. As Table 3-5 shows, Walworth's hospital price jumped from about $3,200 in 1991 to $8,300 in 1994.

The mean of the HMO index we constructed did not change dramatically between 1991 and 1994, moving from 810 to 870, however the distribution about this mean increased

[38] The correlations between the population (as a proxy for sample size) and the standard error of the uninsured estimate were -.666 for 1991 and -.643 for 1994, corroborating the intuition that as sample size increases the accuracy of the estimates tends to increase. The correlations between sample size and population were extremely high at .993 in 1991 and .985 in 1994.

[39] The accuracy of this estimate was double-checked with the Bureau of Labor Statistics.

substantially, with the standard deviation moving from 630 in 1991 to 750 in 1994. This suggests an increased variance across counties in HMO markets as the gap between counties with high and low index values grew. (Note that we converted this variable into a 0-1 dummy based on a natural break in the data.)

There was little change from 1991 to 1994 in the other independent variables. Rounding out the structural vector of variables, the mean of doctors per thousand people did not change, however a number of counties had significant decreases, notably Dane County, dropping from 4.23 to 3.42 physicians per thousand and Milwaukee County which declined to 2.88 from 3.29. None of the firm characteristics measures had changes in their means greater than one percent, and within the counties there was little shift from 1991 to 1994. In the demographic vector, again little change occurred. None of the counties changed their MSA status, however, the nonwhite population increased slightly from 3.3 percent to 3.7 percent, and median income increased from about $30,500 to nearly $36,000.

Empirical Methods

To identify the covariates of county-level uninsurance in Wisconsin, we specified a cross-sectional multivariate regression model and analyzed the two periods separately. Since county-level estimates of the percentage uninsured are aggregations of individual-level data, the equations were estimated using a logistic procedure for a proportional dependent variable constructed from grouped data. Logistic estimation was chosen because the percent uninsured is bounded by 0 and 1 (i.e., 100 percent). Then, the regressions were estimated using weighted least squares, which accounts for heteroscedasticity in grouped data by weighting the dependent variables by a function of the population of the aggregated units (in this case, counties) and the ratio of

uninsurance to one minus the uninsurance rate.[40] Thus, counties with larger sample sizes will be more important in the regression, preventing smaller counties from driving the results.[41]

For 1991 and 1994, then, we estimated:

$$\text{Log (UI/1-UI)} = \alpha + \beta_1 \text{ MARKET} + \beta_2 \text{ ECON} + \beta_3 \text{ DEMOGRAPHICS} + \epsilon$$

where

UI =	Three-year average uninsurance rates, weighted by county population
MARKET =	Full-time equivalent physicians per 1,000 people,[42]
	Physician price index,
	Casemix-adjusted hospital inpatient net revenue per discharge,
	HMO index
ECONOMICS =	Unemployment rate,
	Percent of workers employed in firms with <20 workers,
	Percent of workers earning wages and salaries,
	Percent of workers employed on farms,
	Percent of workers employed in the retail and services sectors,
DEMOGRAPHICS =	Percent nonwhite and Hispanics,
	A dummy for metropolitan areas (MSAs),[43]
	Median income,
α =	a constant;
β_n =	coefficients of the variables; and
ϵ =	an error term, normally distributed with mean zero.

Regression Results. Table 3-6 shows the multivariate results for the 1991 and 1994 county-level uninsurance equations. Results show significant, positive effects of unemployment on uninsurance in both periods, at the 1 percent level in both periods. The effect of unemployment is consistent with our expectations, although it differs from Schmidt and Deichert's (1996)

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[40] This procedure produces minimum chi-squared estimates of the parameters that have been shown to have the same properties as maximum likelihood estimates (Amemiya 1985).

[41] We would not want small counties to drive the results because their uninsurance estimates are potentially less-accurate, owing to their smaller sample sizes.

[42] FTEs are defined according to a 40-hour work week.

[43] To be defined as an MSAs, counties must have one city of at least 50,000 people and an overall population of at least 100,000, according to the Bureau of the Census.

suggestion that unemployment is less important (at least in Nebraska) where it is generally low. In the 1991 period, the marginal effect of unemployment on uninsurance is 0.009. Thus, a one percent increase in unemployment rates would result in an increase in uninsurance of nine tenths of one percentage point. This effect is noticeably smaller than the marginal effect of unemployment in the 1994 period, where a one percent increase in unemployment would yield a 2.3 percentage point increase in uninsurance. This effect seems large but, it might make sense if every person who loses a job also loses insurance for herself and 1.3 family members, on average. It may be that while the marginal job in the tight labor market of 1994 was a source of coverage, perhaps due to the recession the marginal job in 1991 did not offer benefits. We conclude that unemployment has a consistently strong influence on uninsurance, but the size of the effect varies over time.

In addition to the unemployment rate, in the 1991 period, the percentage of workers who are paid wages and salaries is also significant, at the 5 percent level. The marginal effect of the percent wage and salary workers is 0.015, indicating that counties with a one-percent higher percentage of workers receiving wages and salaries (as opposed to proprietors) would have uninsurance levels that are 1.5 percentage points higher.

None of the other variables showed significant effects in the 1991 period. Overall, the model explained 46 percent of the variance in this period, not as much as in the 1994 period. This may be due to problems with the data in the earlier period or to secular events between 1990 and 1992, discussed further below.

The model performed substantially better in the 1994 period, however, with an explained variance of 71 percent. In addition to the unemployment rate, the percentage of workers in the

retail and services sectors and hospital prices also have positive and significant (.01 level) effects on uninsurance rates. Each of these effects is consistent with expectations. The marginal effect of retail and services staff is small: counties with one percent more workers in these sectors would have a five tenths of a percentage point higher uninsurance rate.

To relate hospital price changes directly to uninsurance, we calculated the elasticity of demand for insurance, given hospital inpatient prices.[44] Elasticity in this case measures the change in the fraction of the population with insurance in response to percentage changes in hospital prices. The elasticity of demand for insurance in this period is .997. This indicates an almost one-to-one tradeoff between increases in hospital inpatient prices and decreases in insurance coverage rates. Thus, if hospital inpatient prices in a given county were 10 percent higher than the mean, uninsurance there would be ten percent higher than the statewide average, all other things equal. Recalling the literature review, this elasticity is higher than what other studies have estimated for employees deciding whether to buy insurance, which range from 0 to -0.65 (Marquis and Long 1995; Chernew et al. 1997). Employers may be more price sensitive than employees, however. Our elasticity is lower than other recent estimates for firms, notably Feldman et al. (1997), who concluded that small firms' offer elasticities are between -3.9 and

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[44] Price elasticity of demand is calculated using the formula: $\frac{\%\text{ change in quantity}}{\%\text{ change in price}}$,

or given the form of the equation we estimated, the coefficient of the price multiplied by the price evaluated at the mean:

$$\varepsilon = \beta\bar{x}$$

-5.8.

Physicians per thousand people has significant *negative* effects on uninsurance in the 1994 period (.05 level). The marginal effect of physicians per thousand people is large: in counties with one more physician per thousand people, uninsurance would be 2 percentage points lower (but it should be noted that physician supply does not tend to change much over time). The direction of the effect is consistent with expectations that high physician supply would be associated with higher levels of insurance coverage, either because physician competition reduces prices and may thus reduce the price of insurance or because physicians tend to locate where coverage is good. It may be logical that this effect is stronger in 1994 than in 1991 if price competition among physicians became more intense as managed care plans came to dominate some of the metropolitan markets, like Madison, in Dane County, and the city of Milwaukee. Most of the firm and worker characteristics were insignificant in the 1994 equation. Notably, the percent wage and salary workers is no longer significant in 1994, although it was in 1991, and the percent retail and services workers is significant in 1994, whereas it was not in 1991.

Correlations among the firm and worker variables increases their coefficients' standard errors and may have reduced the significance levels of some of these variables. Table 3-7 shows correlations above r = 0.50. The highest correlation is between the percent workers in small firms and the percent wage and salary workers, which have correlation coefficients of -.880 in 1991 and -.882 in 1994. In 1994, small firms is also correlated with farm workers (r = +.665), and farm workers and wage and salary workers are highly correlated in both years (1991, r =

-.843; 1994, r = -.828). Retail and service workers is not very highly correlated with any of the other firm characteristics.[45]

It is not surprising to learn that these variables are highly correlated with each other in both periods (except for retail and services workers). After all, many small firms are farms, 90 percent of proprietors (the opposite of wage and salary workers) own small firms, and farmers are often proprietors. Several efforts at dropping correlated variables failed to consistently improve the models' fit. Instead, we compared specifications with and without the group of four variables using an F-test. Based on this test, we can conclude that the four variables representing firm or worker characteristics have significant roles in explaining uninsurance in the 1994 model, but we cannot provide estimates of their separate effects.

Other than the hospital price and doctor supply measure, market structure and competition variables were generally insignificant, including the physician price index and the HMO index interaction term. It is interesting that physician prices were not important in our uninsurance models, given the importance of hospital prices. This could be due to the instability in physician pricing patterns resulting from the high profile Marshfield Clinic antitrust case, which was filed between our 2 study periods. Physician and hospital prices are not very highly correlated, so that cannot explain the absence of significance for physician prices. We are confident that the BCBSUW prices we used are representative of typical physician prices because the company's market share, above ten percent in only four counties, is generally too small to set prices in most counties.

[45] It is interesting to note that retail and services workers is correlated with physician supply, perhaps because both are associated with population centers.

The insignificance of the HMO index variable suggests that concentration among HMOs, even where HMO penetration is high, did not substantially affect uninsurance in Wisconsin during the study period. We cannot measure whether or not premium prices are generally higher in such cases, but there is no strong correlation between provider prices and HMO concentration. In any case, premium prices do not appear to change enough as a result of HMO concentration to alter uninsurance rates.

All the demographic variables, including a dummy variable for metropolitan areas, percent nonwhite and median income, were insignificant in the model. Furthermore, there is no reason to think that multicollinearity among these variables or with others has obscured their influence. This suggests that in our model, demographics do not drive uninsurance themselves when economic and market structure variables are accounted for.

Because it was the only variable to be significant in both years, we can conclude with some confidence that unemployment is the most important covariate of county-level uninsurance. Furthermore, firm and worker characteristics also proved to have substantial effects on uninsurance as a group. Finally, there is evidence that the underlying health service market structure (as reflected in hospital prices and physician supply) may affect insurance premiums enough to lead some employers and individuals to forgo or take up insurance coverage.

Limitations. The most serious limitation of this study is that we could only use 29 counties. With limited units of observation, the power of our tests is lower than we would like. Furthermore, even in the 29 counties that were large enough to include, coefficients of variation across the three years in each time period are large, suggesting that the uninsurance estimates vary substantially across years. This throws some doubt on the accuracy of our three-year county-level

uninsurance estimates for both periods. Unfortunately, there are no other county-level estimates available for comparison, so we are confident these are the best estimates available at the present time.

The lackluster results of the 1991 model as compared to the 1994 equation raise several questions. Obviously there is reason for concern when a successful model, applied to a different time period, no longer performs well. It suggests that the model's relationships are not stable over time. This is consistent with Schmidt and Deichert's (1996) experience with the instability of their uninsurance prediction models for Nebraska for 1989 and 1991. The authors found that the accuracy of their prediction model for 1989, when used to predict 1991 uninsurance rates, deteriorated substantially. Schmidt and Deichert felt this implied that "relationships between the uninsured rate and economic indicators at the county level are unstable" (Schmidt and Deichert, 1996:95).

There are a variety of other possible explanations for the weakness of the model in the 1991 period relative to the 1994 period. The recession in the early 1990s could be a factor. Looking at the unemployment variable, for example, some counties were hit harder than others. This larger variance in 1991 imparts noise to the 1991 regression.

There are signs that the data in the earlier period may be less stable than in the later period. There are more high-level correlations among the independent variables in 1991 than in 1994. In addition, 1991 uninsurance rates are correlated with 1994 rates at only a very low level, whereas one would expect a high correlation if the direction and magnitude of change across counties was relatively consistent. Furthermore, we were forced to impute some hospital data for

1991 because of missing values. The imputation process may have affected the overall accuracy of some variables, and therefore, the fit of any model using these variables.

Other potential limitations of this portion of the study are that the results for 29 counties in Wisconsin may not be generalizable to other counties, in Wisconsin or in other states. Furthermore, since we could not use the consecutive years of the survey as a panel data set, we could not use a fixed effects specification, generally preferred methodologically because failing to control for unmeasured effects can lead to mistaken conclusions about the importance of various effects. Thus we must have less confidence in the results of the county model than we have in the state models. Nevertheless, our results do support a reasonable level of confidence that unemployment is the most important predictor of uninsurance at the county level and that both hospital prices and doctor supply can matter on the margin when unemployment is low.

V. MAJOR CONCLUSIONS AND DIRECTIONS FOR FUTURE RESEARCH

The variance of uninsurance is high, but our models proved successful in explaining that variance, especially at the state level. Descriptive data remind us that employment is by far the most important source of health insurance in the U.S. Still, our results suggest that states can influence the performance of their private health insurance markets and affect overall rates of coverage on the margin with careful policy choices. Specifically, guaranteed issue, at least in a package of small group market reforms that includes guaranteed renewal, limits on pre-existing condition restrictions, and portability, appears to increase rates of health insurance coverage at the state level. However, we also learned that states need to be mindful of tradeoffs among policy choices and coverage goals, for the net increase in coverage from issue reforms can be negated if premium rating restrictions are added to the policy mix. Rate restrictions in the small group market would appear to raise more groups' premiums than they lower, probably because of the underlying skewed distribution of health expenditures. More currently insured groups are relatively healthy than not and they enjoy premiums below the market average.

In the nongroup market, all packages of reforms that we tested, including guaranteed issue plus premium rating restrictions, worked to reduce overall coverage. While many fewer states implemented any individual market reforms, and thus the number of cases that triggered the statistical results is necessarily smaller than for small group reforms, the statistical significance of the results was very high. These findings are consistent with the inference that policy interventions in the private individual market, while well-intentioned and likely to help certain high risk individuals, on balance lead more people to drop coverage than are newly able to purchase insurance. Presumably this net decrease in coverage is due to an increase in premiums, which

could reflect that the average risk in the overall nongroup pool increased, as a result of reforms that make it easier for the sick to purchase insurance in a voluntary market. The relatively healthy may decide that insurance is no longer worth the higher price. The net effect of reforms on the insured risk pool in the private nongroup market is an important topic for future study, since one criteria for the success of reforms might be the types of people who end up with coverage compared to those who lose it. Which are healthier, on average? We cannot be sure today.

One important and policy relevant finding was that selective contracting restrictions, at least as implemented in the states through 1995, had no noticeable effect on rates of uninsurance. This suggests that any premium increases from existing state any willing provider and freedom of choice laws were minimal. Most benefit mandates were also insignificant, though required treatment for alcohol and drugs did reduce private coverage and increase overall uninsurance rates. This may be the most expensive mandate benefit mandate we tested.

We found some evidence that the presence of a high risk pool without an enrollment cap reduces Medicaid enrollment and increases private coverage without affecting overall coverage. However, this result was sensitive to our dropping one large state (California) from the analysis, so more work needs to be done to refine our interpretation of this effect. Consistent with this tentative finding, we infer that high risk pools provide an alternative for very sick individuals, and that this opportunity reduces their reliance on Medicaid as well as on the private commercial market. By doing so, high risk pools serve as a vent for bad health risks, freeing nongroup insurers from worry about extreme adverse selection in voluntary insurance markets. While there are not enough high risk pool enrollees to register noticeably in our nation's household surveys, many more individuals may benefit from the effect that high risk pools' existence has on premiums

charged in the individual and even small group markets. More work is justified to understand why this result was both statistically strong in the full model but not robust over all specifications we tested.

At the subnational level, Wisconsin is clearly the current leader in health insurance coverage in the U.S. Yet even among the 29 Wisconsin counties we examined, the variance of uninsurance is rather high, with rates ranging from 4 to 17 percent. The rate of unemployment is the most important predictor of the rate of uninsurance at the county level, though the size of its effect varies over time, we suggest with the business cycle. When the overall economy was strong, the effect of unemployment was large, suggesting that during booms the marginal job comes with health insurance but perhaps does not during recessions.

In the 1994 county model, when the Wisconsin economy was strong, hospital prices and physician supply affected coverage in ways that are consistent with competitive theory. Holding all other things equal, higher hospital prices reduced coverage, presumably through higher premiums. Also, greater physician supply increased coverage, presumably through greater price and quality competition for patients and lower premiums, although it could be that physicians simply prefer to locate where coverage rates are high. The estimated elasticity of hospital prices on rates of uninsurance was -1.0. That is, counties with 10 percent higher hospital prices than average would observe, all else equal, a 10 percent higher than average uninsurance rate. Still, because the county model was necessarily based on relatively few observations, we would like to corroborate them with data from other states and time periods.

One last salient finding from the Wisconsin county study was that the percentage of workers in small firms had no apparent effect on rates of uninsurance. Given national data that

consistently show that small firms are less likely to offer insurance and that workers in small firms are less likely to find coverage in general, this result was surprising and inspiring. A case study of small business and health insurance in Wisconsin might yield useful lessons for other states concerned about the rates at which small firms offer health insurance to their workers.

In conclusion, we note that our major findings from the state model should not be interpreted to imply that all forms of rating restrictions are ill-advised, or that individual market reforms cannot improve insurance coverage. Indeed, the most important direction for further research is in the possibility of varying influences on uninsurance of different forms of rating restrictions. Due to the imprecision of secondary sources, we were unable to calibrate our measures of rating restrictions more precisely than the simple presence or absence of some kind of restriction. It is possible, for example, that looser rate bands, in conjunction with guaranteed issue and like reforms, might be expected to be more conducive to coverage expansions, in conjunction with issue reforms, than are tighter rate bands. Looser bands squeeze existing premiums less severely, and thus may keep a greater proportion of premiums below the threshold level where firms and workers will decline coverage. Similarly, in the individual market, we could only examine the few laws that states have enacted thus far. New formulations of nongroup reforms, especially the creation and expansion of high risk pools pursuant to the federal Health Insurance and Portability Act of 1996, may address the problems of higher risk individuals more efficiently. This general line of research, following from the lessons learned in this study, could be extremely helpful to the policy debate on the effectiveness and future of health insurance market reforms at both the state and federal levels.

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Table 1-1. Demographic Characteristics of the Nonelderly Uninsured in the United States, 1996.

Characteristic		Percentage Uninsured Within Demographic Category
Age	Infants	17.5
	1-5	13.0
	6-12	14.5
	13-17	16.5
	18-20	24.4
	21-24	32.8
	25-34	22.5
	35-44	16.4
	45-54	13.7
	55-64	13.9
Income (Percentage of poverty level)	0-99%	33.8
	100%-124%	35.8
	125%-149%	31.7
	150-199%	27.1
	200%-399%	14.7
	400% or more	7.1
Race	White	13.4
	Black	23.2
	Hispanic	35.1
	Other	21.8
Citizenship	Noncitizen	44.4
	Citizen	15.8
Family Type	Married without Children	16.0
	Married with Children	13.4
	Single without Children	28.0
	Single with Children	20.6

Source: EBRI 1997.

Table 1-2. Regulatory Policies

	SMALL GROUP INSURANCE MARKET				INDIVIDUAL INSU
	Guaranteed Issue, Guaranteed Renewal, Portability and Limits on Pre-Existing Conditions Exclusions	Guaranteed Renewal, Portability and Limits on Pre-Existing Conditions Exclusions	Any other Combination of Guaranteed Issue, Guaranteed Renewal, Portability and Limits on Pre-Existing Conditions Exclusions	Rating Restrictions	Guaranteed Issue and Rating Restrictions
Alabama					
Alaska	93-95			94-95	
Arizona	94-95			94-95	
Arkansas			92-95	92-95	
California	93-95			93-95	
Colorado	94-95		91-93	92-95	
Connecticut	90-95			91-95	
Delaware	93-95			92-95	
Florida	93-95		92	92-95	
Georgia				92-95	
Hawaii					
Idaho	94-95		93	94-95	95
Illinois		94-95		94-95	
Indiana			92-95	93-95	
Iowa	92-95			92-95	
Kansas	93-95		92	92-95	
Kentucky	95			95	95
Louisiana		94-95	93	92-95	
Maine	93-95	92	90-91	93-95	94-95
Maryland	93-95			94-95	
Massachusetts	92-95			92-95	
Michigan					
Minnesota	93-95			93-95	
Mississippi	95			95	
Missouri	94-95		93	93-95	
Montana	94-95			94-95	
Nebraska	94-95		92-93	92-95	
Nevada					
New Hampshire	95		93-94	93-95	95
New Jersey	94-95			93-95	93-95
New Mexico		95	91-94	92-95	
New York	93-95			93-95	93-95
North Carolina	92-95			92-95	
North Dakota	94-95			92-95	
Ohio	93-95			94-95	
Oklahoma	94-95		92-93	93-95	
Oregon	92-95			92-95	
Pennsylvania					
Rhode Island	92-95			93-95	
South Carolina	95	92-94		92-95	
South Dakota		95	92-94	92-95	
Tennessee	94-95	93		93-95	

Table 1-3. More Regulatory Policies and Public Programs

	Any Willing Provider or Freedom of Choice Laws	Mandated Benefit for Alcoholism or Drug Treatment	Mandated Benefit for Mental Health Treatment	Mandated Benefit for Chiropractic Care	High Risk Pool Without an Enrollment Cap	High Risk Pool With an Enrollment Cap
Alabama	89-95	89-95		89-95		
Alaska		89-95		89-95	93-95	
Arizona				89-95		
Arkansas	89-95			89-95		
California			89-95	89-95		91-95
Colorado			89-95	89-95	91-95	
Connecticut		89-95	89-95	89-95	89-95	
Delaware	95			94-95		
Florida	94-95			89-95	89-90	91-95
Georgia	89-95			89-95		
Hawaii		89-95	89-95			
Idaho	95					
Illinois	89-95	89-95		89-95		89-95
Indiana	89-95			89-95	89-95	
Iowa				89-95	89-95	
Kansas		89-95	89-95	89-95	93-95	
Kentucky	95			89-95		
Louisiana	93-95			89-95		92-95
Maine		89-95	89-95	89-95		89-94
Maryland		89-95	89-95	89-95		
Massachusetts	95	89-95	89-95	89-95		
Michigan		89-95		89-95		
Minnesota		89-95	89-95	89-95	89-95	
Mississippi	95	89-95		89-95	92-95	
Missouri		89-91	89-91	89-95	92-95	
Montana	89-93	89-95	89-95	89-95	89-95	
Nebraska		89-95		89-95	89-95	
Nevada	89-95	89-95		89-95		
New Hampshire	89-95		89-95	89-95		
New Jersey	95	89-95		89-95		
New Mexico				89-95	89-95	
New York		89-95		89-95		
North Carolina	94-95			89-95		
North Dakota	90-95	89-95	89-95	89-95	89-95	
Ohio		89-95	89-95	89-95		
Oklahoma	94-95			89-95		
Oregon		89-95	89-95		90-95	
Pennsylvania		89-95		89-95		
Rhode Island		89-95		89-95		
South Carolina	95			89-95	90-95	
South Dakota	89-95			89-95		
Tennessee				89-95	89-95	
Texas	89-95	89-95		89-95		
Utah	89-95	89-95		89-95	91-95	
Vermont		89-95				
Virginia	89-95	89-93	89-95	89-95		

Table 1-4. State Low Income Population and Medicaid Eligibility and Enrollment, 1994-1995

State	% Population <100% FPG	% Population Eligible for Medicaid	% Population Enrolled in Medicaid
United States	18.9	19.4	15.8
Alabama	22.8	14.7	13.2
Alaska	14.0	16.3	14.9
Arizona	20.2	21.1	16.6
Arkansas	19.4	17.3	13.3
California	22.8	26.0	21.9
Colorado	11.1	9.2	8.8
Connecticut	14.1	17.4	11.4
Delaware	13.6	12.9	11.9
District of Columbia	33.1	27.7	22.2
Florida	21.3	19.4	16.2
Georgia	19.4	19.6	17.0
Hawaii	20.2	25.1	18.6
Idaho	16.5	15.6	11.2
Illinois	17.6	20.2	16.4
Indiana	15.2	12.9	10.9
Iowa	14.8	15.5	11.4
Kansas	16.8	17.2	11.8
Kentucky	22.9	22.3	17.9
Louisiana	29.0	25.5	18.1
Maine	12.5	17.3	15.6
Maryland	15.9	17.6	11.8
Massachusetts	14.9	15.7	13.2
Michigan	17.4	19.1	15.6
Minnesota	12.7	15.2	10.8
Mississippi	29.2	22.7	20.0
Missouri	17.7	15.7	14.7
Montana	17.7	17.3	12.8
Nebraska	11.5	12.2	10.3
Nevada	15.7	12.8	9.7
New Hampshire	9.4	18.6	8.5
New Jersey	13.7	13.5	10.9
New Mexico	29.7	22.8	18.9
New York	21.8	22.6	18.3
North Carolina	18.0	17.0	15.4
North Dakota	12.6	11.6	9.1
Ohio	16.7	14.4	14.3
Oklahoma	20.4	18.7	15.1
Oregon	16.8	29.6	15.7
Pennsylvania	16.8	17.3	14.7
Rhode Island	14.9	20.0	13.8
South Carolina	20.7	18.8	13.4
South Dakota	17.2	13.6	10.8
Tennessee	19.9	34.0	26.2
Texas	22.9	17.8	15.2
Utah	11.4	15.1	11.0
Vermont	10.0	24.5	16.1
Virginia	14.8	14.5	10.8
Washington	16.4	23.5	16.3
West Virginia	23.1	25.4	24.0
Wisconsin	12.3	14.3	11.3
Wyoming	13.2	10.5	9.7

Table 2-1. Means and Standard Deviation of the Model Variables

	Mean	Standard Deviation
Dependent Variables		
Percent uninsured	15.0%	4.3%
Percent with private coverage	73.0%	6.5%
Percent Medicaid	9.8%	3.1%
Packages of Small Group Reform		
Guarant'd issue, guarant'd renewal, portability and limits on prexisting exclusions	27.1%	44.5%
All of the above except guaranteed issue	4.9%	21.5%
Any other combination of the above mentioned small group reforms	10.0%	30.0%
Small Group Rating Restrictions		
Any rating restriction	40.0%	49.1%
Packages of Individual Insurance Market Reforms		
Guaranteed issue and rating restrictions	4.6%	20.9%
Any other combination of guarant'd issue, guarant'd renewal, portability, limits on prexisting exclusions, and rating restrictions	4.3%	20.3%
Other Reforms		
AWP or FOC laws- strong or medium versions	30.3%	46.0%
Mandated benefit for alcoholism or drug treatment	58.3%	49.4%
Mandated benefit for mental health treatment	34.9%	47.7%
Mandated benefit for chiropractic care	82.6%	38.0%
Public Programs		
High risk pool without enrollment cap	34.3%	47.5%
High risk pool with enrollment cap	7.7%	26.7%
Percent eligible for Medicaid	15.5%	4.6%
Market Structure and Competition		
HMO share of insured population	16.7%	12.4%

Table 2-2. Values of Dependent Variables for 1989 and 1995

	Percent Uninsured		Percent Private		Percent Medicaid	
State	1989	1995	1989	1995	1989	1995
United States	**15.0%**	**15.5%**	**74.3%**	**70.4%**	**8.7%**	**12.2%**
Alabama	19.2	15.4	70.6	72.1	7.4	10.9
Alaska	23.1	11.9	66.6	71.2	6.8	11.1
Arizona	20.2	20.4	70.5	62.7	6.4	12.8
Arkansas	18.9	20.7	69.1	66.0	9.1	10.7
California	20.1	19.6	65.8	61.4	12.4	17.9
Colorado	16.5	15.1	75.3	78.3	6.2	5.1
Connecticut	9.3	9.3	86.5	81.8	3.3	7.4
Delaware	16.8	16.5	74.1	75.1	5.9	7.8
Florida	20.2	19.9	69.9	64.6	7.0	13.2
Georgia	16.8	16.4	73.0	67.6	8.3	14.0
Hawaii	9.2	4.2	80.4	76.9	7.7	16.3
Idaho	17.0	16.4	77.0	74.2	4.4	8.1
Illinois	10.4	10.8	78.1	76.0	10.1	12.0
Indiana	14.4	12.9	78.2	79.0	5.2	6.9
Iowa	7.7	12.5	85.3	79.8	6.0	7.6
Kansas	10.3	14.2	81.5	74.7	7.4	9.0
Kentucky	12.4	15.5	74.9	68.7	11.1	13.2
Louisiana	19.1	22.0	65.4	59.3	12.9	16.7
Maine	10.5	12.1	80.0	75.4	7.8	10.6
Maryland	11.1	16.2	80.2	72.6	7.1	9.8
Massachusetts	9.5	12.2	80.6	77.9	8.0	8.7
Michigan	9.5	9.5	77.6	78.6	11.7	11.2
Minnesota	9.8	8.4	81.8	80.6	7.2	7.8
Mississippi	18.7	20.5	62.8	61.3	15.0	16.6
Missouri	12.9	14.7	76.8	74.0	8.0	9.6
Montana	16.9	14.4	72.9	69.8	7.8	10.2
Nebraska	11.7	9.6	79.8	80.2	6.9	7.5
Nevada	17.9	19.9	75.2	71.9	4.0	6.7
New Hampshire	16.3	11.1	79.1	81.0	2.8	6.3
New Jersey	11.4	15.3	81.1	76.3	6.7	7.6
New Mexico	24.8	28.4	64.5	50.5	7.1	16.6
New York	13.2	16.4	73.8	67.7	11.5	14.7
North Carolina	15.4	15.1	75.1	71.1	6.9	11.6
North Dakota	9.9	9.6	83.9	81.3	4.3	6.5
Ohio	8.7	13.6	81.2	75.3	8.5	10.1
Oklahoma	20.9	20.4	67.0	64.3	8.2	12.1
Oregon	14.3	15.1	76.9	73.2	7.1	9.7
Pennsylvania	9.8	10.8	80.9	77.0	7.8	11.1
Rhode Island	11.9	12.7	79.8	74.8	7.4	10.7
South Carolina	15.7	16.0	73.6	68.2	7.6	11.8
South Dakota	15.6	11.0	76.4	76.4	5.0	6.7
Tennessee	14.4	8.8	72.7	68.0	10.1	21.5
Texas	24.2	24.4	65.7	61.5	7.2	12.5
Utah	9.7	11.7	82.9	80.5	5.8	6.9
Vermont	9.5	12.2	82.8	74.4	6.4	11.5
Virginia	14.5	14.3	77.2	74.5	6.0	7.8
Washington	11.9	12.4	77.0	73.2	8.5	11.8
West Virginia	14.1	16.5	70.8	65.3	12.8	17.0
Wisconsin	9.9	7.6	82.7	81.6	6.3	9.1
Wyoming	13.6	17.3	78.5	72.2	5.2	8.5

Table 2-3. Values of Selected Independent Variables for 1989 and 1995

	HMO Share of Insured Population		Hospital expenses per adjusted patient day		Percent Non-white and Hispanic	
State	1989	1995	1989	1995	1989	1995
Alabama	6.3%	10.7%	547.43	819.25	27.4%	35.2%
Alaska	0.0%	0.0%	995.90	1,340.84	25.8%	25.5%
Arizona	25.5%	40.9%	805.33	1,191.47	30.4%	37.9%
Arkansas	3.4%	21.9%	495.86	704.06	18.6%	19.9%
California	40.7%	56.6%	871.80	1,315.27	44.3%	51.3%
Colorado	28.9%	33.7%	689.35	1,069.23	19.9%	18.2%
Connecticut	26.1%	38.7%	763.37	1,263.88	17.3%	20.0%
Delaware	24.9%	40.3%	712.61	1,057.76	21.6%	26.4%
Florida	16.2%	34.7%	717.60	1,004.12	29.4%	36.7%
Georgia	6.4%	12.8%	577.08	836.11	30.6%	36.3%
Hawaii	31.0%	28.6%	549.90	955.79	67.7%	77.2%
Idaho	2.7%	5.2%	494.72	718.57	8.3%	11.8%
Illinois	16.7%	25.4%	665.06	1,049.76	26.6%	29.0%
Indiana	9.9%	13.4%	614.78	962.92	10.9%	8.9%
Iowa	9.8%	6.2%	456.27	702.04	4.4%	6.5%
Kansas	11.3%	8.9%	476.70	732.02	12.3%	13.9%
Kentucky	10.5%	21.5%	520.45	794.79	8.6%	9.7%
Louisiana	8.1%	16.1%	658.12	902.02	35.0%	33.1%
Maine	2.6%	12.8%	523.66	916.41	2.1%	1.5%
Maryland	20.3%	41.7%	623.75	1,064.10	31.5%	36.6%
Massachusetts	31.7%	50.7%	729.80	1,156.73	12.8%	15.5%
Michigan	19.1%	27.9%	675.12	993.76	18.3%	18.3%
Minnesota	32.7%	35.1%	506.76	736.30	6.6%	8.4%
Mississippi	0.0%	1.7%	413.30	583.83	37.8%	42.6%
Missouri	13.4%	33.3%	624.03	966.80	13.8%	15.5%
Montana	0.5%	4.1%	378.21	493.30	8.8%	12.5%
Nebraska	7.3%	13.8%	441.60	661.17	8.0%	9.0%
Nevada	12.1%	26.4%	840.88	1,071.80	22.1%	29.8%
New Hampshire	13.9%	28.5%	613.74	915.06	2.8%	3.1%
New Jersey	15.1%	31.6%	567.92	962.17	27.6%	29.4%
New Mexico	19.2%	22.4%	682.80	1,073.30	51.2%	52.3%
New York	19.4%	40.1%	581.74	908.57	32.4%	36.2%
North Carolina	7.1%	16.1%	549.44	831.91	25.8%	29.3%
North Dakota	4.2%	1.7%	396.18	521.49	6.3%	5.4%
Ohio	15.2%	24.3%	668.75	1,061.20	13.4%	14.8%
Oklahoma	7.9%	15.6%	585.74	861.04	19.8%	20.4%
Oregon	30.0%	59.6%	742.55	1,141.04	9.9%	11.7%
Pennsylvania	13.5%	36.2%	629.13	963.09	13.0%	15.1%
Rhode Island	30.1%	33.7%	620.79	1,098.16	11.0%	14.1%
South Carolina	3.0%	11.9%	533.60	923.46	32.3%	36.6%
South Dakota	5.7%	3.7%	396.08	476.05	9.6%	7.5%
Tennessee	6.7%	16.8%	584.47	871.25	18.0%	21.5%
Texas	10.9%	18.4%	685.08	1,062.84	40.6%	46.5%
Utah	16.5%	37.8%	773.65	1,212.73	9.0%	12.6%
Vermont	6.0%	16.8%	556.41	713.56	2.0%	2.2%
Virginia	8.8%	12.6%	588.74	901.22	24.5%	28.0%
Washington	18.0%	30.4%	746.14	1,318.47	13.9%	15.3%
West Virginia	5.0%	10.1%	533.83	762.65	4.2%	2.6%
Wisconsin	29.4%	32.0%	523.00	794.37	9.4%	10.8%
Wyoming	0.7%	0.0%	432.43	545.10	9.3%	8.7%

Table 2-4. Results for State-level Analysis

		Dependent Variable				
		UNINS		PRIVATE		MEDICAID
Packages of Small Group Reform	N	Coefficient	Marginal Effect	Coefficient	Marginal Effect	Coefficient
Guarant'd issue, guarant'd renewal, portability and limits on prexisting exclusions	95	-0.1107***	-0.0141	0.0448*	0.0088	0.0172
All of the above except guaranteed issue	17	-0.0673*	-0.0086	0.0335		0.0170
Any other combination of the above mentioned small group reforms	35	-0.0768*	-0.0098	0.0131		0.0412
Small Group Rating Restrictions						
Any rating restriction	140	0.1072***	0.0136	-0.0546**	-0.0107	-0.0072
Packages of Individual Insurance Market Reforms						
Guaranteed issue and rating restrictions	16	0.1062***	0.0135	-0.0501*	-0.0099	-0.0161
Any other combination of guarant'd issue, guarant'd renewal, portability, limits on prexisting exclusions, and rating restrictions	15	0.0814**	0.0103	-0.0673***	-0.0132	-0.0056
Other Reforms						
AWP or FOC laws- strong or medium versions	106	-0.0203		0.0211		0.0298
Mandated benefit for alcoholism or drug treatment	204	0.1506*	0.0191	-0.1288**	-0.0254	0.0397
Mandated benefit for mental health treatment	122	-0.1075		0.1380		-0.1572
Mandated benefit for chiropractic care	289	0.1588		-0.0700		0.0052
Public Programs						
High risk pool without enrollment cap	120	-0.0161		0.0596**	0.0117	-0.1453***
High risk pool with enrollment cap	27	-0.0520		0.0399		-0.0141
Percent eligible for Medicaid	350	-0.0092**	-0.0012	-0.0108***	-0.0021	0.0263***
Market Structure and Competition						
HMO share of insured population	350	0.0078***	0.0010	-0.0046**	-0.0009	0.0004
Number of physicians per 100,000 people	350	-0.0014		0.0008		-0.0012
Number of hospital beds per 100,000 people	350	0.0002		-0.0001		-0.0003
Hospital expenses per adjusted patient day	350	0.0001		0.0001		-0.0002
Employment and Economic Conditions						
Unemployment rate	350	-0.0053		0.0003		0.0045

Table 2-5. High Correlations among Independent Variables, State Analysis

Variables	Correlation	
	Correlation Variables	r=
Small group insurance market- guaranteed issue, guaranteed renewal, portability and limits on pre-existing conditions exclusions	Small group rating restrictions	+.67
HMO share of the insured population	Physicians per 100,000 people	+.66
	Adjusted inpatient expenditures per day	+.59
	Percent college graduates	+.53
	Median income in thousands	+.51
	Percent nonmetropolitan	-.59
Physicians per 100,000 people	Percent college graduates	+.67
	Median income in thousands	+.63
Hospital expenses per adjusted patient day	Median income in thousands	+.56
	Hospital beds per 100,000 people	-.64
Percent workers working full-time	Percent nonwhite and Hispanic	+.52
	Percent workers in a small firm	-.55
Percent workers in a small firm	Percent nonmetropolitan	+.53
	Percent married	+.53
Median income in thousands	Percent college graduates	+.73

Table 3-1. Sample Size and Confidence Intervals for the 1991 and 1994 Periods			
Range of observations in a county	Number of counties	Mean Standard Error	Mean width of the 95% confidence interval
100-149	14	2.97	11.6
150-199	10	2.66	10.1
200-249	13	1.92	7.6
250 and over	38	1.38	5.4

Table 3-2. Type of Service Categories and Procedures Included in Physician Price Index

Category, Code and Procedure	Category, Code and Procedure
Primary care	**Surgery**
99203 Office Visit, New Patient, 30 minutes	43235 Upper Gastrointestinal Endoscopy
99213 Office Visit, Established Patient, 15 minutes	58120 Dilation and Curettage
99214 Office Visit, Established Patient, 30 minutes	58150 Total Hysterectomy
99244 Office Consult, New or Established Patient, 60 minutes	66984 Cataract removal with Lens Implant
90843 Psychiatric Visit, 20-30 minutes	69436 Tympanostomy, General Anesthesia
90844 Psychiatric Visit, 45-50 minutes	69440 Middle Ear Exploration
93000 Electrocardiogram	69450 Tympanolysis, Transcanal
Hospital Visits	**Imaging**
99222 Initial Hospital Care, New or Established Patient, 50 minutes	70450 Computerized Axial Tomography Scan, Head or Brain
99254 Initial Inpatient Consultation, 80 minutes	71020 X-Ray, Chest, Two Views
	76805 Echography, Pregnant Uterus
Obstetric Care	**Laboratory Tests**
59400 Total Obstetric Care, Vaginal Delivery	81000 Urinalysis, Routine
59410 Vaginal Delivery only	87081 Culture, Bacterial, Screening Only
59510 Total Obstetric Care, Cesarean Delivery	88305 Surgical Pathology
59515 Cesarean Delivery and Postpartum Care	

Source: Norton 1995.

Table 3-3. Descriptive Statistics, County Analysis.

County model	**1991**				**1994**			
	Mean	**Std. Dev**	**Min.**	**Max.**	**Mean**	**Std. Dev**	**Min.**	**Max.**
Uninsured Rate	10.6	3.2	5.1	19.4	9.9	3.6	3.6	17.0
Percent of employees in firms with under 20 employees	27.0	4.3	20.0	36.4	26.6	4.3	19.4	36.4
Unemployment rate	5.4	1.9	2.3	12.5	4.6	1.2	2.3	7.2
Doctors per 1000 people	1.6	1.2	0.6	5.6	1.6	1.1	0.5	5.3
Case mix adjusted hospital price/1000	3.7	0.8	2.4	6.1	5.0	1.3	3.1	8.3
Physician cost index	0.9	0.1	0.8	1.2	1.0	0.1	0.9	1.1
HMO index dummy variable	0.3	0.5	0.0	1.0	0.3	0.5	0.0	1.0
Percent of employees working on farms	4.7	3.5	0.0	14.6	4.1	3.2	0.0	14.1
Percent of employees deriving income from wages and salaries	83.4	4.1	73.4	91.4	83.6	4.2	72.6	91.5
Percent of employees in the retail and services sectors	47.7	5.8	39.2	63.4	47.8	6.0	38.9	67.0
Percent of population which is not white	3.3	4.5	0.5	23.7	3.7	4.9	0.6	25.8
Dummy variable for MSAs	0.6	0.5	0.0	1.0	0.6	0.5	0.0	1.0
Median income/ 1000	30.5	5.0	22.4	44.6	35.9	5.7	27.6	51.5

Table 3-4. County Uninsured Rates and 95% Confidence Intervals

County	1990-1992				1993-1995			
	Uninsured Rate	95% Confidence Interval	1992 Population	Sample Size	Uninsured Rate	95% Confidence Interval	1994 Population	Sample Size
Brown	10.5	± 2.3	198,602	668	8.8	± 2.0	207,269	777
Chippewa	12.7	± 4.5	52,929	207	12.1	± 5.0	54,007	162
Columbia	12.1	± 4.2	45,955	237	8.0	± 3.8	48,374	201
Dane	8.9	± 1.3	374,713	1,762	6.5	± 1.2	390,261	1,491
Dodge	10.2	± 3.1	77,395	372	8.6	± 3.0	78,264	327
Eau Claire	12.6	± 4.0	86,113	267	13.9	± 3.7	87,939	333
Fond du Lac	9.1	± 3.1	90,878	339	10.0	± 3.2	92,951	329
Grant	9.3	± 4.0	49,195	203	13.2	± 4.7	49,705	197
Jefferson	7.9	± 3.2	68,478	268	13.6	± 3.8	72,405	304
Kenosha	10.5	± 2.8	131,624	467	11.0	± 3.2	137,810	370
La Crosse	11.5	± 3.5	98,645	317	6.8	± 3.0	101,004	266
Manitowoc	13.6	± 3.8	80,858	307	11.3	± 3.7	82,145	287
Marathon	7.2	± 2.5	116,870	407	11.1	± 2.8	120,111	480
Marinette	13.3	± 5.1	40,893	169	17.0	± 6.0	41,846	152
Milwaukee	13.7	± 1.1	956,869	3,758	13.0	± 1.0	938,105	4,757
Outagamie	11.6	± 2.6	141,997	590	4.5	± 1.7	147,458	556
Ozaukee	5.1	± 3.0	74,469	205	3.6	± 2.4	78,026	237
Portage	14.0	± 4.5	62,349	227	14.8	± 4.6	64,040	229
Racine	13.1	± 2.5	177,594	710	10.4	± 2.7	181,704	495
Rock	13.3	± 2.6	141,222	657	9.4	± 2.6	145,958	481
St. Croix	5.2	± 3.0	51,290	206	15.4	± 5.0	53,994	199
Sauk	19.4	± 4.9	47,687	254	6.6	± 3.4	50,234	200
Sheboygan	9.9	± 3.1	104,479	359	6.7	± 2.5	107,031	373
Walworth	7.1	± 3.2	76,319	243	13.4	± 4.5	80,720	221
Washington	7.7	± 2.8	98,550	354	4.1	± 2.3	107,234	275
Waukesha	10.0	± 1.8	312,954	1,058	5.3	± 1.4	332,207	995
Waupaca	14.6	± 5.1	46,844	182	10.6	± 4.2	48,468	209
Winnebago	6.8	± 2.1	142,880	554	7.8	± 2.2	147,869	565
Wood	7.2	± 3.3	74,350	243	9.9	± 3.4	75,702	290

Table 3-5. 29 County Values for Selected Variables

	Uninsured Rate		Unemployment Rate		Doctors per 1000		Hospital Price		HMO Index		Physcian Index	
	1991	1994	1991	1994	1991	1994	1991	1994	1991	1994	1991	1994
Brown	0.10	0.09	4.80	4.30	1.58	1.65	4.42	5.54	637.55	25.11	1.01	0.95
Chippewa	0.13	0.12	6.20	6.10	0.96	1.08	3.25	3.80	547.04	654.51	0.89	0.91
Columbia	0.12	0.08	5.90	6.90	0.84	0.84	2.42	3.09	1115.30	1370.85	0.91	0.92
Dane	0.09	0.06	2.40	2.30	4.23	3.42	4.68	7.28	1529.91	1648.64	0.97	1.06
Dodge	0.10	0.09	5.80	4.40	0.83	0.85	3.15	4.11	546.90	1212.90	0.91	0.88
Eau Claire	0.13	0.14	5.70	4.60	2.03	2.37	3.68	3.52	1358.73	1537.34	0.84	0.96
Fond du Lac	0.09	0.10	4.90	4.00	1.11	1.41	3.93	4.13	69.58	102.60	0.97	0.87
Grant	0.09	0.13	5.40	5.30	0.58	0.52	2.45	3.42	938.71	784.51	0.94	0.92
Jefferson	0.08	0.14	5.50	4.30	0.96	0.90	3.19	4.52	1307.69	851.06	0.92	0.97
Kenosha	0.10	0.11	6.10	5.00	1.00	1.11	3.16	5.87	736.06	589.75	1.01	1.05
La Crosse	0.12	0.07	4.30	4.10	3.55	3.73	3.74	4.68	617.14	539.25	0.95	0.95
Manitowoc	0.14	0.11	5.60	4.90	1.01	1.13	3.19	4.86	9.16	11.29	0.93	0.93
Marathon	0.07	0.11	5.40	5.40	1.68	1.70	4.69	6.51	1995.00	2156.10	0.86	1.11
Marinette	0.13	0.17	9.10	7.20	0.95	0.91	2.91	4.64	74.23	0.92	0.84	0.90
Milwaukee	0.14	0.13	5.50	5.20	3.29	2.88	4.87	7.06	1082.99	1187.03	1.06	1.04
Outagamie	0.12	0.05	2.30	2.60	1.59	1.81	4.31	6.18	953.83	1844.56	0.94	0.95
Ozaukee	0.05	0.04	3.90	3.10	0.96	1.18	3.80	4.49	708.85	764.48	1.05	1.03
Portage	0.14	0.15	4.10	5.30	1.12	1.21	4.01	5.28	48.61	73.84	0.92	0.94
Racine	0.13	0.10	6.70	5.80	1.14	1.26	4.56	4.49	532.41	462.43	1.16	1.09
Rock	0.13	0.09	12.50	5.20	1.45	1.41	3.19	4.47	562.34	801.14	0.90	0.88
St. Croix	0.05	0.15	5.70	3.80	1.33	0.54	3.08	5.34	386.45	492.14	0.87	0.91
Sauk	0.19	0.07	6.50	4.90	1.28	1.32	3.70	4.18	1488.51	1298.33	0.95	0.94
Sheboygan	0.10	0.07	5.70	3.30	0.81	1.28	3.55	4.37	29.43	133.56	0.93	0.88
Walworth	0.07	0.13	3.70	2.90	0.79	0.87	3.17	8.32	633.52	696.73	0.89	0.94
Washington	0.08	0.04	4.90	3.80	0.91	0.75	2.96	4.49	689.54	653.08	1.01	1.06
Waukesha	0.10	0.05	4.20	3.60	1.24	1.52	3.94	3.69	720.80	735.27	1.07	1.02
Waupaca	0.15	0.11	5.40	5.60	0.87	0.71	2.68	3.85	575.93	640.03	0.93	0.92
Winnebago	0.07	0.08	3.70	3.90	1.63	1.78	3.57	6.46	542.59	1150.32	0.95	0.91
Wood	0.07	0.10	4.60	5.00	5.62	5.28	6.06	6.90	2920.05	3482.41	0.94	1.04

Table 3-6. Results of County Level Uninsurance Analysis

County Model	1991		1994	
	Coeffcient	Marginal Effect	Coeffcient	Marginal Effect
Percent of employees in firms with under 20 employees	0.056		-0.019	
Unemployment rate	0.095 ***	0.009	0.255 ***	0.023
Doctors per 1000 people	0.038		-0.224 **	-0.020
Case mix adjusted hospital price/1000	-0.212		0.199 ***	0.018
Physician cost index	1.816		-0.733	
HMO index dummy variable	0.252		-0.138	
Percent of employees working on farms	0.089		0.067	
Percent of employees deriving income from wages and salaries	0.162 **	0.015	0.030	
Percent of employees in the retail and services sectors	-0.006		0.052 ***	0.005
Percent of population which is not white	-0.017		-0.005	
Dummy variable for MSAs	-0.015		-0.222	
Median income/ 1000	-0.006		0.013	
Intercept	-18.626 **	-1.768	-8.348 *	-0.745

Adjusted R squared, 1991=.464, 1994 =.710
Prob. > F, 1991 = .021, 1994 = 0.000
N = 29

* = Significant at the 10% level, ** = Significant at the 5% level, *** = Significant at the 1% level

Elasticity for 1994 case mix adjusted price= 0.997

Table 3-7. High Correlations among Independent Variables, County Analysis

Variable	Correlations above .5 between Independent Variables			
	1991		1994	
	Correlated Variables	r =	Correlated Variables	r =
Percent of workers in firms with under 20 workers (Below20)	Hosprice Wageslry	-.671 -.880	Farmwork Wageslry	+.665 -.882
Doctors per capita (Doccap)	Hosprice Retlserv	+.752 +.518	Wageslry Retlserv	+.611 +.546
Casemix adjusted hospital price (Hosprice)	Farmwork Wageslry Below20 Doccap	-.534 +.699 -.671 +.752		
Percent of workers on farms (Farmwork)	Hosprice Docindex Wageslry MSA	-.534 -.512 -.843 -.564	Below20 Wageslry	+.665 -.828
Unemployment rate (Ue_rate)			Medinc	-.615
Percent of workers paid by wage or salary (Wageslry)	Below20 Hosprice Farmwork Perntwht	-.671 +.699 -.843 +.563	Below20 Doccap Farmwork Perntwht	-.882 +.611 -.828 -.558
Percent of workers in the retail and services industries (Retlserv)	Doccap	+.518	Doccap	+.546
Percent of population which is not white (Perntwht)	Wageslry	+.563	Wageslry	+.558
MSA dummy variable (MSA)	Farmwork	-.564		
Median income (Medinc)	Docindex	+.520	Ue_rate	-.615
Physician cost index (Docindex)	Medinc	+.520		