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How Did SCHIP Affect the Insurance Coverage of Immigrant Children?

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Abstract

The State Children's Health Insurance Program (SCHIP) significantly expanded public insurance eligibility and coverage for children in "working poor" families. Despite this success, it is estimated that over 6 million children who are eligible for public insurance remain uninsured. An important first step for designing strategies to increase enrollment of eligible but uninsured children is to determine how the take-up of public coverage varies within the population. Because of their low rates of insurance coverage and unique enrollment barriers, children of immigrants are an especially important group to consider. In this paper, we compare the effect of SCHIP eligibility on the insurance coverage of children of foreign-born and native-born parents. In contrast to results of research on the earlier Medicaid expansions, we find similar take-up rates for the two groups. Also in contrast to earlier work, we find evidence of lower crowd-out among immigrants. This suggests that state outreach strategies were not only effective at increasing take-up overall, but were successful in reducing disparities in access to coverage.

JEL Classification: I1

Keywords: Medicaid, SCHIP, Health Insurance, Crowd-Out, Take-Up, Immigrants

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I. Introduction

In the past two decades there have been substantial initiatives at the state and federal levels aimed at increasing insurance coverage among children. Most recently, the State Children's Health Insurance Program (SCHIP) expanded public insurance eligibility for children in "working poor" families. SCHIP significantly increased public insurance coverage and decreased the rate of uninsured among children in families with incomes between 100 and 300 percent of the federal poverty level (Cunningham, Hadley and Reschovsky 2002; Lo Sasso and Buchmueller 2004; Hudson, Selden and Banthin 2005). By 2002, roughly half of all children in the US were income-eligible for some kind of public health insurance (Selden, Hudson and Banthin 2004). However, despite this success, it is estimated that over 6 million children who are eligible for public insurance remain uninsured. These children represent a majority of all uninsured children. Extending coverage to these eligible but uninsured children is an important but challenging objective for federal and state policy makers.

A crucial first step for addressing this problem is to determine how the take-up of public coverage varies within the population. Children of immigrants are an especially important group to consider. Previous research shows that foreign-born adults are nearly three times as likely to be uninsured as native-born Americans (Buchmueller, Lo Sasso, Lurie and Dolfin, in press) and that children of immigrants are also more likely to be uninsured than children whose parents were born in the US (Ku and Matani 2001). Immigrants' lower rate of insurance coverage is driven mainly by a lower rate of employer-sponsored insurance, which in turn is largely explained by differences in human capital and the types of jobs held by immigrants and native-born workers. While this makes public insurance more important as a source of coverage for

children of immigrants, because of language and cultural barriers they may be less likely than children in non-immigrant families to enroll.

Despite the well documented gap in insurance coverage, there has been surprisingly little research on how public insurance take-up differs between immigrants and natives. One study of the Medicaid expansions found a weaker response to Medicaid eligibility among children of foreign-born parents as compared to children whose parents were born in the US (Currie 1999). Because SCHIP was enacted just after the 1996 Federal welfare reform legislation, which singled out recent immigrants for less generous benefits, there is additional reason to expect a lower take-up of SCHIP among the children of immigrants. On the other hand, in implementation of SCHIP included much greater emphasis on outreach including marketing campaigns in languages other than English (Aizer 2004). If these efforts were effective, they may have reduced nativity-related differences in take-up. Moreover, because non-natives are so much less likely to have private insurance than natives, it is possible that the problem of "crowd-out"—i.e., the substitution of public insurance for private coverage—may be less of an issue for immigrants. However, recent research on the impact of welfare reform found that reductions in public coverage among immigrants were completely offset by increases in private coverage, a striking finding implying 100% substitution of private coverage for public coverage (Borjas 2003).

In this paper we test whether the effect of the SCHIP expansion was different for children of foreign-born and US-born parents. The analysis is based on repeat cross-section data from the Current Population Survey and the same research design employed successfully in previous research on the effects of Medicaid and SCHIP expansions on insurance coverage for the entire population of children (Cutler and Gruber 1996; Lo Sasso and Buchmueller 2004). Specifically,

we use an instrumental variables approach in which the effect of SCHIP eligibility is identified by cross-state differences in the timing and extent of changes in the income eligibility limit over the period from 1996 to 2001. We test for the effect of SCHIP on insurance coverage from any source as well as on the probability of having public insurance (take-up) and on the probability of having private coverage (crowd-out).

In contrast to earlier studies, our results suggest that take-up among the children of the foreign born was at least as high as natives. Estimates of the effect of eligibility on reported coverage by any private insurance suggest that there was little crowd-out for either group. However, earlier research on SCHIP suggests that some survey respondents misclassify public insurance provided through private carriers as private, non-group coverage. When we estimate separate models for non-group and group insurance, the results suggest that the increased SCHIP enrollment of children of immigrants coincides with a decline in employer-sponsored coverage. In contrast, we find little evidence of crowd-out for children of native-born parents.

II. Background and Previous Literature

The SCHIP Program

SCHIP was established by Federal legislation in 1997 and enacted by states in the next several years. Like prior studies, we exploit variation in the timing and extent of the SCHIP eligibility expansions to identify effects on coverage.[1] Eleven states implemented their program in 1997, 34 did so in 1998 and by 2000 every state had a program in place. Variation in the extent of expansion comes from differences across states in income thresholds both before and after implementation. Prior to SCHIP, states were required to cover children 6 years of age and

[1] Details on when states implemented SCHIP and how income eligibility limits changed are provided in Appendix Table A-1.

under up to 133 percent of poverty, though they were allowed to expand coverage up to 185 percent and still receive federal matching dollars.[2] Because there were no such Federal standards for older children, there was much more variation in eligibility limits. Since the implementation of SCHIP, in most states income eligibility limits are the same for children of all ages. In 2000, the last year of our data, the modal income eligibility threshold was 200% of the FPL (18 states). Nine states expanded eligibility even further and the other 26 states had income limits of between 133% and 192% of the FPL.

Because states were given considerable flexibility, state programs vary in other dimensions as well (Wolfe and Scrivner 2005; Bansak and Raphael 2006). States were allowed to experiment with different strategies for disseminating information about the program, simplifying the application and enrollment process and improving retention. These efforts may explain why SCHIP appears to have had a stronger effect on public insurance coverage than earlier Medicaid expansions that were targeted at children in families with incomes above the poverty line. LoSasso and Buchmueller (2004), Hudson, Selden and Banthin (2005) and Bansak and Raphael (2006) find that 8 to 10% of children who gained income eligibility for SCHIP enrolled in the program. This is comparable to the take-up rate that Card and Shore-Sheppard (2004) estimate for an earlier Medicaid expansion target at children in families with incomes up to 100% of the FPL and larger than what they find for expansions affecting children with family incomes between 100% and 133% of the FPL.

Other differences across states have to do with program design and rules. States were given the option of expanding their Medicaid program, establishing a new stand-alone program, or both. States also varied in the strategies used to limit crowd-out. The most common approach

[2] As of 1996 several states (CA, MN, RI, TN, VT and WA) had used state funds to expand eligibility for some children even further.

is to require that children must be without insurance for some period prior to enrolling. Thirty-three states have such waiting periods, ranging from three to twelve months. These differences in program features raise issues for estimating the effect of SCHIP on insurance coverage. For the purposes of the present study, it is important to consider not only whether specific design features are likely to affect take-up or crowd-out, but also whether the effect is likely to vary with nativity. For example, prior studies indicate that mandatory waiting periods are effective at reducing crowd-out (LoSasso and Buchmueller 2004; Kronebusch and Elbel 2004; Bansak and Raphael 2006). Because of differences in access to private coverage, waiting periods are less likely to be binding for children of immigrants than for those of native-born parents. Therefore, failing to account for waiting periods may understate take-up rates for natives relative to immigrants. The evidence on differences between Medicaid expansions and new stand-alone programs is less clear. LoSasso and Buchmueller find no significant difference between the two types of programs, but Bansak and Raphael find that public coverage increased more in states that chose to expand Medicaid than in states that introduced a new SCHIP program. Whatever explains these divergent results, there are not strong *a priori* reasons to expect a differential effect of program type by nativity.

Immigrants and Public Health Insurance

The existing study that is most similar to ours is one by Currie (1999) that compares the response of children of immigrants and children with native-born parents to Medicaid expansions occurring between 1989 and 1992. Currie finds that increased eligibility led to higher Medicaid enrollment among children of native-born parents, but had no significant effect for children of immigrants. Medicaid eligibility did reduce private insurance coverage for immigrant children,

however. She interprets this pattern as indicating that the Medicaid expansions induced some immigrant parents to drop private coverage in favor of the "conditional coverage" for emergency care to which they were entitled even if they did not formally enroll in the program. Families that drop private coverage when they become eligible for Medicaid reap a financial benefit by forgoing monthly premium contributions while maintaining the ability to receive free care in the case of emergency. However, conditional coverage is not likely to improve access to primary or preventive health care. Indeed, while Currie finds that increased program eligibility led to greater use of ambulatory care for children of native-born parents, she finds no effect for children of immigrants.

SCHIP was passed just after landmark federal welfare reform legislation, the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA). In addition to setting time limits for cash welfare payments, PRWORA restricted the eligibility of immigrants for welfare and other public programs, including Medicaid. Under Federal law, immigrants arriving in the US after 1996 are prohibited from receiving Medicaid for five years. However, states have the option of using their own funds to insure new immigrants and a number have done so. Initially, the legislation also restricted eligibility for immigrants arriving prior to 1996, though those provisions were never enacted. Nonetheless, some analysts argue that by creating confusion about eligibility rules and contributing to fears of deportation, PRWORA had a "chilling effect", decreasing program participation among immigrants who remained eligible for these programs. Several studies show that since the enactment of PRWORA, Medicaid enrollment has fallen considerably and that the decline is greater for foreign-born compared to native-born persons (Borjas 2003; Kandula et al. 2004; Kaushal and Kaestner 2005).

III. Data

Our data are drawn from the March Current Population Survey (CPS) for the years of 1997 to 2001, which provide information on household health insurance coverage for the period from 1996 to 2000.[3] Since most states enacted SCHIP in 1998, this provides between two to three years of data prior to and after the enactment of SCHIP. These five years of data provide a sample size of 181,402 children who are less than 18 years old, living with their parents and not heads of their own households. Because parental nativity is a key variable in our analysis, we exclude observations for which this information is missing, giving us a sample of 167,298. Because the SCHIP expansions should not have had any effect on the public insurance eligibility or coverage of higher income children, in addition to analyzing the full sample of children, we conduct all analyses on a "target" sample consisting of children in families who are at or below 300% of the FPL (N=109,059).

A key methodological issue for this analysis concerns the way children are categorized according to their parents' nativity. In our main analysis, we follow Borjas (2003) in categorizing children based on the nativity of the head of their household. In the full sample, we have 130,689 children in families headed by a native-born adult (hereafter "natives") and 36,609 children in families where the household head is foreign-born (hereafter "non-natives"). In the targeted sample, there are 83,943 natives and 25,116 non-natives. As we describe below, we obtain similar results when we define non-native children as those with at least one foreign-born parent. Table 1 provides summary statistics for native and non-native children in both the full

[3] Although the questionnaire refers to health insurance coverage during the prior calendar year, some research suggests that many respondents do in fact report current coverage status (Swartz 1986; Berger, Black and Scott 1998). Nonetheless, following the previous literature, we interpret the insurance variables as referring to the previous year. All the years of data we use are after a change in the insurance questions that occurred in 1995 (Swartz 1997).

and targeted samples.

IV. Estimation Strategy

We use the repeated cross-section data from the CPS to estimate several versions of the following regression model:

$$COVERAGE_i = \alpha PUBELIG_i + \beta \boldsymbol{X_i} + \gamma_s \boldsymbol{STATE_i} + \gamma_t \boldsymbol{TIME_i} + \varepsilon_{1i}, \quad (1)$$

where the dependent variable $COVERAGE_i$ represents the type of health insurance held by child *i*: public, private, or uninsured. *PUBELIG* is an indicator for public insurance eligibility, which is constructed based on the child's age, family income and the eligibility standards effective in the child's state of residence at that time. The vector $\boldsymbol{X}$ includes the child's age and standard socio-demographic characteristics. We include a full set of year and state dummies to account for national trends in health insurance coverage and long-standing differences across states. The equation is estimated as a linear probability model.

All models are estimated on samples that are stratified by nativity. Given the link between employment and health insurance coverage in the US, our baseline specification includes several variables to account for the possibility that the state and year dummies do not fully capture changes in labor market opportunities for different subpopulations. We interact the year dummies with categorical variables for education and race to account for the fact that workers in different "skill" groups may have been affected differently by changes in macroeconomic conditions over this period. To account for state-specific economic shocks, we also include several regressors that vary by state and year: the state-level unemployment rate, the

gross state product (GSP) and the percentage of the state's population receiving cash welfare benefits each year. The unemployment rate and GSP are included to account for the relationship between local area macroeconomic conditions and insurance that has been documented by prior studies (Cawley and Simon 2003; Glied and Jack 2003). The average caseload is intended to capture cross-state differences in the effect of welfare reform. We also estimate models that replace these state-level variables with full state/year interactions. This specification has the advantage of accounting for possible state-specific macroeconomic shocks in the most flexible way. However, it demands a lot of the data, leaving little residual variation for identifying the effect of SCHIP.

As noted, most states restrict SCHIP eligibility to children who have been without private insurance for a certain period of time. Because this eligibility criterion is based on one of our outcome variables, we cannot incorporate it directly in the construction of *PUBELIG*. Therefore, our regression estimates will understate the marginal take-up rate among children meeting all eligibility criteria. A rough adjustment can be made by dividing the coefficient on *PUBELIG* by the percent of children in the sample who were uninsured. In addition, because there is variation across states in the length of the waiting period, we can estimate the effect of this policy parameter on coverage by augmenting the regression model as follows:

$$COVERAGE_{ci} = \alpha_{3c} PUBELIG_i + \alpha_{4c} PUBELIG_i \times MONTHS + \beta_c X_i + \gamma_c STATE_i + \theta_c TIME_i + \varepsilon_{ci}. \quad (2)$$

In this equation, *MONTHS* equals the number of months a child who meets the program's income eligibility standards must be uninsured before qualifying for SCHIP. We expect the

length of the waiting period to have a negative effect on public coverage. If the waiting period was effective at reducing crowd-out, the effect on private coverage should be positive.

To account for the endogeneity of public insurance eligibility, we use the same instrumental variables strategy as previous studies on eligibility expansions (Currie and Gruber 1996a; Cutler and Gruber 1996; Ham and Shore-Sheppard 2001; Lo Sasso and Buchmueller 2004; Hudson, Selden and Banthin 2005). Specifically, we instrument for *PUBELIG* using a simulated eligibility measure generated by applying the eligibility rules for each state in each year to a nationally representative sample of children. The instrument is the mean imputed eligibility for each state-year-age combination. Because our model includes state and year fixed effects, identification comes from variation within states in the timing of SCHIP implementation and the extent to which SCHIP raised income eligibility limits. Additional within-state variation comes from the fact that the magnitude of the eligibility changes differed by child age.

All of this variation in eligibility affects children in families with incomes less than 300% of the FPL. Fitting these models to a sample of all children assumes that in the absence of SCHIP, trends in insurance coverage for children in the SCHIP "target group" would have been similar to children in higher income families, who remained ineligible for public insurance. Our results are subject to bias if this assumption does not hold. Therefore, in addition to conducting the analysis on the full sample, we estimate all models on a subsample of children in families with incomes below 300% of the FPL.

V. Results

Trends in Public Insurance Eligibility and Insurance Coverage

Before turning to the regression results, we present unadjusted trends in public insurance eligibility and insurance coverage in Table 2. In addition to reporting data for the full samples of native and non-native children, we report results for children in families with incomes below the poverty line and those with family incomes between 100% and 300% of the FPL.

The full sample results (panel A) show that in 1996, non-native children were significantly more likely to be eligible for public health insurance: 44% vs. 27%. Eligibility increased in the next three years. It leveled off for natives between 1999 and 2000 and fell for non-natives. The latter result is likely caused by a combination of changes in family income and sampling error as no states restricted eligibility between 1999 and 2000. By 2000, 39% of native children and 58% of non-natives were eligible for either Medicaid or a stand-alone SCHIP program. However, for both groups, actual public coverage actually fell between 1996 and 1998, before increasing by the end of the period. Private insurance coverage increased by roughly 4 percentage points for both groups causing the percentage without insurance to decline slightly. Uninsurance declined for both groups over the period; falling by nearly 5 percentage points for non-natives and by 2 percentage points for natives.

The data for children in families with incomes below the poverty line (Panel B) show that that the decline in public insurance coverage between 1996 and 1998 was a result of welfare reform, as documented in prior studies (Garrett and Holohan 2000; Kaestner and Kaushal 2003; Cawley, Schroeder and Simon 2006). The data on eligibility indicate that these children were not directly affected by the SCHIP eligibility expansions: in 1996 over 90% were already eligible for Medicaid. It has been suggested, however, that Medicaid enrollment of already eligible

children increased as a result of SCHIP outreach efforts. This may explain why the percentage of poor children with public insurance increased after 1998. Private insurance increased for both groups, with a slightly larger change for natives. The net effect was a reduction in the percent uninsured of between 1 and 2 percentage points.

The increase in public insurance eligibility over this period was concentrated among children with family incomes between 100 and 300% of the FPL (Panel C). In 1996, only 13% of children in this income range were eligible for Medicaid. By 2000, nearly half of native children and almost three-fifths of non-natives were eligible. For both natives and non-natives, actual public enrollment fell in 1997 but increased thereafter. The growth in enrollment was stronger for non-natives (a change of 10 percentage points) than for natives (3 percentage points). These unadjusted figures imply that the marginal take-up rate was roughly three times as large for non-natives compared to natives: 29.5% vs. 9.5% percent. However, recall that our eligibility measure does not account for the fact that in most states children who already had private insurance were not eligible for SCHIP. We can do a rough adjustment by dividing these figures by the percentage of children without private insurance. Doing the adjustment based on the 1996 values implies that 43% of native children (0.095/[1 - 0.776]) and 68% of non-native children (0.295/[1- 0.570]) who met income eligibility requirements and did not have private insurance took up SCHIP coverage. There was essentially no change in private coverage among native children in this income group and a decline of 1.9 percentage points for non-natives.

Regression Results

Regression results corresponding to equation (1) are reported in Table 3. The results in the top panel are for all children, while in the bottom panel the sample is limited to children in

families with incomes less than 300% of the FPL. For each sample, results from our baseline model are reported in columns 1 (natives) and 2 (non-natives) and results from a model with state-year interactions are reported in columns 3 (natives) and 4 (non-natives). State-year interaction terms allow our model to account for very general forms of policy endogeneity, enabling us to have greater confidence that the parameter estimates better reflect the underlying change in Medicaid/SCHIP policy. However, because there is less policy variation *within state-year* we anticipate larger standard errors. Because of this trade-off we choose to present both sets of estimates.

In the first row of each panel the dependent variable equals one for children who are reported to have public insurance coverage and zero for those without public coverage. For all samples and specifications, the coefficient on our eligibility variable is positive and statistically significant. In the full sample, the models without the state-year interactions the coefficient for children of foreign-born parents is roughly 60 percent larger than the coefficient children whose parents were born in the US (.113 vs. .070). Adding state-year interactions increases the native coefficient slightly (to .080) and reduces the non-native coefficient slightly (to .110). Because the variation in eligibility comes almost entirely from changes in rules affecting children in lower income families, a comparison of the full sample results and the results for children with family incomes below 300% of the FPL provides a specification check. The fact that the parameter estimates are so similar suggests that our results are not being driven by other factors that affected the insurance coverage of higher income children.

For all models and all samples, we find no statistically significant effect of public eligibility on private coverage. The results in column 2 are suggestive a small crowd-out effect for non-natives, with a coefficient of -0.037 in the full sample and -0.033 in the low income

sample. However, each of these coefficient estimates has a t-statistic of roughly 1. Adding state-year interactions reduces the size of each coefficient. The results for natives are even less suggestive of crowd-out. As a result, when the dependent variable equals one for children who are uninsured (bottom panel), the estimated coefficients on the eligibility variable are negative and of a comparable magnitude as the positive coefficients in the take-up model.

Importantly, the uninsurance coefficients are uniformly smaller (in absolute value) than the public insurance coefficients, suggesting that increases in public coverage are not accounted for on a one-to-one basis with decreases in uninsurance, which is suggestive of crowd-out. The differences between the public and uninsurance coefficients are generally larger for natives than for non-natives, indicating potentially larger crowd-out for natives.

Table 4 presents results from estimates of equation (2), in which the eligibility variable is interacted with the number of months the state requires a child to be uninsured before enrolling in SCHIP. (For convenience we only display results for the children under 300% of FPL; results for all children are very similar.) The expectation is that the coefficient on the length of the waiting period should be negative in the take-up regression and positive in the private insurance regression. This is, in fact, what we find for both natives and non-natives. This indicates that waiting periods have the intended effect of reducing crowd-out and, by achieving this goal, reducing public coverage. In particular, each additional month in a state's waiting period reduces take-up by roughly 2 percentage points for both the native and non-native groups. The results also imply that if states had no waiting period, take-up rates would range from 9-13 for natives and 15-18 for non-natives, which is considerably higher than the take-up rates observed in Table 3.

Table 5 presents results that split private coverage into private non-group and private group insurance. Private non-group arguably should not be affected by the SCHIP expansions—indeed, one might expect non-group coverage to fall as more people with non-group insurance become eligible for SCHIP. However, prior work has found that non-group coverage appears to increase with the implementation of SCHIP (Lo Sasso and Buchmueller 2004), suggesting that respondents in the CPS might not understand that their coverage is provided through Medicaid/SCHIP via a private carrier or with a new name that does not evoke a state-sponsored program. Premium requirements for SCHIP in some states might also lead to confusion in responding to CPS insurance coverage questions. Such misinterpretations might be even more common among non-natives as they may lack familiarity with public programs. The CPS is clearer in asking about private group coverage, specifying the need for an employer or union's sponsorship of the insurance coverage.

The results in Table 5 indicate that for natives there are significant decreases in private group coverage, but only in the models with state-year fixed effects. However, we do not observe statistically significant increases in private non-group insurance for natives. Although for non-natives there is some evidence of increases in reports of non-group coverage coincident with increases in public eligibility; the result is statistically significant in the regression without state-year fixed effects, but insignificant in the model including state-year fixed effects. Results for group coverage for non-natives are quite variable between specifications and do not suggest consistent results, perhaps owing to smaller sample sizes.

Discussion

Our analysis suggests important differences in the response to SCHIP between children of natives and non-natives. First, we find that take-up rates are at least as high for non-natives as they are for natives. This is an important result because it implies that outreach efforts embedded in the SCHIP expansion may have led to greater success in enrolling non-native children. It is also at odds with the previously mentioned study of the earlier Medicaid expansions by Currie (1999), which suggested lower take-up rates for immigrant children.

Second, we find evidence consistent with previous research that suggests some degree of misclassification of public insurance coverage in CPS data, but the effect appears to be largely driven by non-natives. This is perhaps not entirely surprising given that non-native families are likely to have less familiarity with state programs and the US health insurance system overall.

Third, we find that crowd-out of private health insurance appears to be more prevalent among natives. Lower crowd-out among non-natives is at odds with the findings of Borjas (2003) whose results imply that all non-natives who lost Medicaid coverage acquired private coverage in response. This result suggested a 100% degree of substitution between public and private coverage for immigrants. Assuming substitution is symmetric between losing and gaining public coverage, our results suggest crowd-out is not as significant an issue as suggested by Borjas (2003). Combining coefficients across Tables 3 and 5, crowd-out for natives ranges from 65-80% while for non-natives crowd-out ranges from 0-37%.[4]

Fourth, the impacts of waiting periods as anti-crowd-out mechanisms have generally similar effects for both natives and non-natives. For both groups additional months of duration for the waiting period lead to the intended effect of reducing private insurance substitution and the unintended effect of reducing take-up.

[4] We use the <300% FPL sample with state-year fixed effects for our back of the envelope crowd-out estimates.

Taken together our results suggest that the SCHIP expansions of the late 1990s were largely successful in increasing enrollment of the children of non-natives. This finding suggests that SCHIP was a comparative success story with respect to its outreach efforts. However, more work is needed to elucidate what specific aspects of the outreach efforts were most efficacious in order to inform policy makers should future expansions be contemplated.

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Table 1. Summary Statistics

	All Children		Children Below 300%FPL	
	Natives	**Non-Natives**	**Natives**	**Non-Natives**
Uninsured	0.106	0.276	0.147	0.320
Public Insurance	0.187	0.263	0.285	0.317
Private Insurance	0.733	0.488	0.605	0.393
Public Eligibility	0.335	0.538	0.537	0.669
Number of Person in HH	4.237	4.774	4.294	4.894
Two-Parent Household	0.710	0.780	0.587	0.742
Male	0.511	0.513	0.511	0.511
White	0.730	0.152	0.650	0.103
Hispanic	0.115	0.663	0.148	0.746
Black	0.126	0.066	0.168	0.066
Other Non-white	0.034	0.145	0.040	0.112
0 Workers in household	0.102	0.113	0.158	0.137
1 Worker in household	0.395	0.438	0.475	0.477
2+ Workers in household	0.503	0.449	0.367	0.386
1+ Worker in a large firm	0.681	0.554	0.604	0.500
0 Adults with some college	0.360	0.517	0.489	0.689
1 Adult with some college	0.299	0.204	0.319	0.197
2+ Adults with some college	0.341	0.204	0.192	0.114
Total # with Fair/Poor Health	0.155	0.221	0.201	0.248
MSA Residence	0.699	0.919	0.643	0.909
Child is Native-Born	0.997	0.780	0.997	0.764
Child is Foreign-Born	0.003	0.220	0.003	0.236
AFDC/TANF caseload	0.012	0.015	0.012	0.015
State Unemp Rate	4.582	5.045	4.624	5.088
Sample Size	138,468	28,823	85,907	23,148

Table 2: Descriptive Trends in Health Insurance Coverage for Children of Natives and Non-Natives, 1996-2000

A. All Children

Native	1996	1997	1998	1999	2000	Change 1996 to 2000
Public Eligibility	0.274	0.281	0.352	0.389	0.387	0.113
Public Coverage	0.203	0.186	0.176	0.180	0.187	-0.016
Private Coverage	0.716	0.720	0.733	0.745	0.754	0.038
Uninsured	0.108	0.116	0.117	0.102	0.085	-0.023
Non-Native						
Eligibility	0.436	0.427	0.605	0.630	0.580	0.144
Public	0.269	0.248	0.255	0.263	0.280	0.011
Private	0.470	0.487	0.482	0.488	0.513	0.042
Uninsurance	0.293	0.292	0.288	0.269	0.241	-0.052

B. Family Income less than FPL

Native	1996	1997	1998	1999	2000	Change 1996 to 2000
Eligibility	0.943	0.967	0.998	1.000	1.000	0.057
Public	0.616	0.580	0.549	0.561	0.580	-0.036
Private	0.260	0.265	0.301	0.309	0.309	0.050
Uninsurance	0.182	0.203	0.216	0.197	0.171	-0.011
Non-Native						
Eligibility	0.939	0.952	0.999	1.000	1.000	0.061
Public	0.497	0.492	0.458	0.482	0.481	-0.016
Private	0.171	0.191	0.194	0.195	0.200	0.029
Uninsurance	0.365	0.352	0.368	0.345	0.346	-0.019

C. Family Income 100% to 300% of FPL

Native	1996	1997	1998	1999	2000	Change 1996 to 2000
Eligibility	0.135	0.149	0.333	0.460	0.480	0.346
Public	0.126	0.114	0.116	0.143	0.159	0.033
Private	0.776	0.773	0.777	0.766	0.766	-0.009
Uninsurance	0.124	0.134	0.129	0.117	0.102	-0.022
Non-Native						
Eligibility	0.138	0.153	0.550	0.628	0.574	0.436
Public	0.150	0.134	0.179	0.203	0.252	0.102
Private	0.570	0.571	0.541	0.538	0.551	-0.019
Uninsurance	0.317	0.325	0.311	0.281	0.242	-0.075

Table 3: Summary OLS Regression Results of Health Insurance Coverage for Children

	Full Sample of Children			
	Natives	**Non-Natives**	**Natives**	**Non-Natives**
Public Insurance				
Mean	0.187	0.263	0.187	0.263
Eligible for Public Insurance	0.070**	0.113**	0.080**	0.110**
	(0.015)	(0.030)	(0.028)	(0.049)
R^2	0.324	0.234	0.329	0.245
Private Insurance				
Mean	0.733	0.488	0.733	0.488
Eligible for Public Insurance	0.008	-0.037	-0.016	-0.006
	(0.018)	(0.033)	(0.032)	(0.053)
R^2	0.308	0.299	0.318	0.301
Uninsurance				
Mean	0.106	0.276	0.106	0.276
Eligible for Public Insurance	-0.069**	-0.074**	-0.048*	-0.103*
	(0.015)	(0.033)	(0.026)	(0.054)
R^2	0.044	0.086	0.055	0.093
State x Year Interaction	N	N	Y	Y
Number of Observations	138,468	28,823	138,468	28,823

	Children Below 300% FPL			
	Natives	**Non-Natives**	**Natives**	**Non-Natives**
Public Insurance				
Mean	0.285	0.317	0.285	0.317
Eligible for Public Insurance	0.063**	0.100**	0.069**	0.107**
	(0.014)	(0.030)	(0.025)	(0.047)
R^2	0.283	0.200	0.289	0.216
Private Insurance				
Mean	0.605	0.393	0.605	0.393
Eligible for Public Insurance	-0.012	-0.033	-0.030	-0.015
	(0.016)	(0.032)	(0.028)	(0.049)
R^2	0.260	0.211	0.269	0.224
Uninsurance				
Mean	0.147	0.320	0.147	0.320
Eligible for Public Insurance	-0.050**	-0.078**	-0.026	-0.083*
	(0.014)	(0.033)	(0.024)	(0.050)
R^2	0.043	0.066	0.055	0.084
State x Year Interaction	N	N	Y	Y
Number of Observations	85,907	23,148	85,907	23,148

Table 4: Summary OLS Regression Results of Health Insurance Coverage for Children below 300% FPL, Interactions with Enrollment Waiting Period

	Natives	Non-Natives	Natives	Non-Natives
Public Insurance				
Mean	0.284	0.319	0.284	0.319
Eligible for Public Insurance	0.133**	0.186**	0.092**	0.150**
	(0.019)	(0.040)	(0.028)	(0.055)
Eligible x Waiting Period	-0.021**	-0.020**	-0.020**	-0.018**
(months)	(0.002)	(0.004)	(0.003)	(0.005)
R^2	0.298	0.204	0.295	0.217
Private Insurance				
Mean	0.607	0.390	0.607	0.390
Eligible for Public Insurance	-0.055**	-0.083*	-0.055	-0.026
	(0.021)	(0.042)	(0.031)	(0.057)
Eligible x Waiting Period	0.014**	0.018**	0.015**	0.015**
(months)	(0.002)	(0.004)	(0.003)	(0.005)
R^2	0.270	0.222	0.274	0.224
Uninsurance				
Mean	0.145	0.322	0.145	0.322
Eligible for Public Insurance	-0.071**	-0.105*	-0.023	-0.114
	(0.018)	(0.044)	(0.026)	(0.059)
Eligible x Waiting Period	0.006**	0.000	0.003	0.000
(months)	(0.002)	(0.004)	(0.002)	(0.005)
R^2	0.035	0.056	0.054	0.073
State x Year Interaction	N	N	Y	Y
Number of Observations	85,907	23,148	85,907	23,148

Table 5: Summary OLS Regression Results of Private Health Insurance Coverage for Children below 300% FPL, Interactions with Enrollment Waiting Period

	Natives	Non-Natives	Natives	Non-Natives	Natives	Non-Natives	Natives	Non-Natives
Private Non-Group Insurance								
Mean	0.092	0.044	0.092	0.044	0.090	0.487	0.090	0.487
	0.001	0.038**			0.005	0.025		
Eligible for Public Insurance	(0.011)	(0.015)			(0.019)	(0.025)		
			-0.003	0.050*			0.003	0.027
Eligible for Public Insurance			(0.014)	(0.020)			(0.021)	(0.029)
Eligible x Waiting Period			0.001	-0.003			0.002	-0.001
(months)			(0.001)	(0.002)			(0.002)	(0.003)
R2	0.060	0.028	0.059	0.023	0.063	0.067	0.063	0.067
Private Group Insurance								
Mean	0.516	0.347	0.516	0.347	0.054	0.451	0.054	0.451
	-0.013	-0.058			-0.059**	0.027		
Eligible for Public Insurance	(0.016)	(0.030)			(0.029)	(0.047)		
			-0.052*	-0.134**			-0.082***	0.024
Eligible for Public Insurance			(0.022)	(0.040)			(0.032)	(0.054)
Eligible x Waiting Period			0.012**	0.021**			0.017***	0.002
(months)			(0.002)	(0.004)			(0.003)	(0.005)
R2	0.277	0.221	0.289	0.236	0.290	0.261	0.297	0.262
State x Year Interactions	N	N	N	N	Y	Y	Y	Y
Number of Observations	85,907	23,148	85,907	23,148	85,907	23,148	85,907	23,148