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# HOW DID SCHIP AFFECT THE INSURANCE COVERAGE OF IMMIGRANT CHILDREN?

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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. We compare the effect of SCHIP eligibility on the insurance coverage of children of foreign-born and native-born parents. In contrast to research on the earlier Medicaid expansions, we find similar take-up rates for the two groups. 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.

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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 and growing group to consider. One in five children under age 18 is either an immigrant or is a member of an immigrant family; since 1990 the number of children in immigrant families has risen seven times faster than the number in native families (Morse 2000). Previous research shows that foreign-born adults are nearly three times as likely to be uninsured as native-born Americans (Buchmueller et al 2007) 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 of the late 1980s and early 1990s 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, the SCHIP legislation included much greater emphasis on outreach, including marketing campaigns in languages other than English (Aizer 2003, forthcoming). If these efforts were effective, they may have reduced nativity-related differences in take-up.

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 on health insurance 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 (CPS) 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 (Currie and Gruber 1996; Cutler and Gruber 1996; Ham and Shore-Sheppard 2001; Lo Sasso and Buchmueller 2004; Hudson, Selden and Banthin 2005). 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 2000. 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 research, 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 work on SCHIP suggests that some survey respondents misclassify public insurance provided through private carriers as private, non-group coverage. Correcting for this misclassification tended to increase take-up estimates and increase crowd-out estimates for both groups. The range of crowd-out estimates was generally similar between natives and non-natives.

## II. Background and Previous Literature

### ***The SCHIP Program***

SCHIP was established by Federal legislation in 1997 and enacted by states over 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. Cross-state variation

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

in the extent of expansion comes from differences in income thresholds both before and after implementation. Prior to SCHIP, states were required to cover children 6 years of age and 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 2007). States were given the option of expanding their Medicaid program, establishing a new stand-alone program, or both. The SCHIP legislation required states to implement mechanisms to limit the crowding out of private insurance, but did not prescribe a specific approach. The most common strategy that states have adopted 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. Some prior studies indicate that these mandatory waiting periods have been effective at reducing the substitution of public insurance for private coverage (LoSasso and Buchmueller 2004; Kronebusch and Elbel 2004; Bansak and Raphael 2007), though a recent study by Gruber and Simon (2006) find that they had no effect.

Compared to the earlier Medicaid expansions the SCHIP legislation placed a greater emphasis on and provided more funding for outreach efforts. States were allowed to experiment

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

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 (2007) find that 8 to 10% of children who gained income eligibility for SCHIP enrolled in the program. While this take-up rate may seem low, it is comparable to the 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. Moreover, because these SCHIP take-up estimates do not account for the fact that many children who meet income eligibility rules are not actually eligible for SCHIP because they already have private coverage, they understate take-up among children meeting all eligibility requirements (Cunningham 2003).

***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 shortly 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, and 14 states used their own funds to provide Medicaid benefits to recent immigrants (Tumlin et al. 1999). 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 (Fix and Passel 1999, 2002; Borjas 2001, 2003; Kaushal and Kaestner 2005; Lurie 2007). 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).

While welfare reform may have depressed public insurance enrollment of immigrants and their children relative to natives, some of the strategies that states used to increase SCHIP take-up are likely to have worked in the opposite direction. It is widely believed that the complexity of Medicaid eligibility rules and administrative procedures was a barrier to enrollment that limited the impact of earlier eligibility expansions. Some of these administrative reforms were specifically targeted at immigrant families. For example, parents are not required to provide proof of their own citizenship when applying for coverage for their children and procedures for verifying the legal status of the children were greatly simplified. Other administrative reforms, such as the use of shorter application forms and reductions in the amount of documentation required, while not directly targeted at immigrants, may have been especially beneficial to families with little prior experience with the system and limited English proficiency.

States with large immigrant populations tailored their outreach efforts to meet the needs of those populations. Several states use community-based organizations to identify eligible families and to assist them with the application process. Aizer (2003, forthcoming) studies the effect of such organizations in California. She finds that proximity to an organization providing bilingual application assistance increased public insurance enrollment by up to 46 percent for Hispanic and Asian children. She also finds a significant effect of Spanish language television ads on the enrollment of Hispanic children.

## 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

[3] There are several well-known limitations of the CPS data on health insurance. Although the questionnaire refers to 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

three years of data prior to the enactment of SCHIP and two to three years after. After 2001, income eligibility limits largely remained constant, but some states responded to the recession by limiting SCHIP and Medicaid enrollment in other ways, while other states pursued parental expansions using SCHIP funds (Aizer and Grogger 2003; Busch and Duchovny 2005). Because modeling these alternative expansion strategies adds considerable complexity and creates greater potential policy endogeneity problems we opt to examine the initial program implementation by studying the period through 2000. 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.

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"). We obtain similar results when we define non-native children as those with at least one foreign-born parent.[4] Table 1 provides summary statistics for native and non-native children. As has been previously documented, non-natives have considerably higher uninsurance rates and considerably lower private insurance coverage rates, regardless of whether we restrict the sample to those in families at or below 300% of the federal poverty level. Note also that the vast majority of children in immigrant families are themselves native-born citizens.

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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).

[4] These results are not reported, but are available upon request.

## IV. Estimation Strategy

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

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

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.[5] The vector ***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 using a linear probability models for each insurance type.

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

[5] Lo Sasso and Buchmueller (2004) estimate models that allow for differences between states that implemented SCHIP by expanding their existing Medicaid programs and those that established new stand-alone programs. Because those models indicated no statistically significant differences between these two approaches, we use a single measure of eligibility for public insurance.

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:

$$COVERAGE_{ci} = \alpha_{3c}\, PUBELIG_i + \alpha_{4c}\, MONTHS_i + \boldsymbol{\beta}_c \boldsymbol{X}_i + \boldsymbol{\gamma}_c \boldsymbol{STATE}_i + \boldsymbol{\theta}_c \boldsymbol{TIME}_i + \varepsilon_{ci}. \qquad (2)$$

In this equation, *MONTHS* represents the number of months a child who meets the program's income eligibility standards must be uninsured before qualifying for SCHIP. In the pre-SCHIP period this variable will take the value of zero for all children. In the post-SCHIP period, it will be zero for children in states without any waiting periods and for children eligible for Medicaid rather than SCHIP; it will be non-zero for children eligible for an SCHIP related expansion program for states that employed waiting periods.[6] 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.

During the time period we analyze, many states contracted with private insurers to provide coverage to SCHIP and Medicaid beneficiaries; others designed their programs to resemble private insurance in order to reduce stigma and thereby make coverage more attractive. These strategies might make it difficult to distinguish public and private insurance in surveys like the CPS as many parents whose children are covered through Medicaid or SCHIP may report that coverage as private non-group. Such classification errors would explain the finding that increases in public insurance eligibility caused by the implementation of SCHIP was associated with an increase in the percentage of CPS respondents reporting that their children had private non-group coverage (Lo Sasso and Buchmueller 2004). Cantor et al. (2007) provide additional evidence that supports this interpretation.

There is less reason to expect public insurance to be reported as group coverage because explicit mention to an employer or union is made in the questionnaire. Therefore, in addition to estimating models that assume the reported insurance variables as accurate, we estimate models

[6] In order to assign a value of MONTHS to children, it is necessary to infer eligibility based on income in state non-Medicaid expansion programs. Hence there is an implicit public eligibility calculation in creating MONTHS, leading to the likely endogeneity of the variable. We instrument for MONTHS using the interaction of simulated eligibility (described below) with the state's waiting period.

in which the dependent variable equals one if either public insurance or non-group private insurance is reported. This composite variable should capture all increases in public insurance coverage, whether they be accurately reported or misreported as increases in non-group private coverage.

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 1996; 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 the variation in eligibility affects children in families with incomes less than 300% of the FPL. Fitting the 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 will be subject to bias if this assumption does not hold. However, selecting on income introduces the potential for endogenous sample selection bias. Therefore, while we present findings for the sample of children in families with incomes below 300% of the FPL, we also estimate the same models for the full sample of children. The results for the full sample (which we do not report, but are available upon request) are very similar to the lower income sample results, lending credibility to our identification strategy.

## 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 latter group can be viewed as the target of the SCHIP expansions.

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, 58% of non-native children and 39% of 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 modestly 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.

The data for children in families with incomes below the poverty line (Panel B) suggest that that the decline in public insurance coverage between 1996 and 1998 was potentially a result of welfare reform, as documented in prior studies (Garrett and Holohan 2000; Kaestner and Kaushal 2003; Bitler, Gelbach and Hoynes 2005; Cawley, Schroeder and Simon 2006). The data on eligibility indicate that for the most part 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 groups 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*

The first set of regression results are reported in Table 3. Results from our baseline model (Equation 1) are reported in columns 1 (natives) and 2 (non-natives). The results in columns 3 and 4 are from models in which we include state and year interaction terms. State-year interactions 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 cells defined by state and year, we anticipate larger standard errors. Because of this trade-off we choose to present both sets of estimates.

In panel A the dependent variable equals one for children who are reported to have public insurance coverage and zero for those without public coverage. For both groups and both specifications, the coefficient on our eligibility variable is positive and statistically significant. For natives, the estimated coefficient on *PUBELIG* is .07 in the baseline model and .08 in the model with state-year interaction terms. Contrary to what might have been expected based on Currie's (2000) results, eligibility for public insurance has a *larger* effect on the public insurance coverage of non-native children. The *PUBELIG* coefficient is .103 in column 2 and .123 in column 4. We find a negative but statistically insignificant effect of public eligibility on the probability that a child is reported to be covered by private health insurance (Panel B). The effect appears somewhat more negative for non-natives, though as with natives the coefficient is not significantly different from zero. As a result, when the dependent variable equals one for children who are uninsured (Panel C), the estimated coefficients on the eligibility variable are negative and of a comparable magnitude as the positive coefficients in the take-up model.

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. 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% to 13% for natives and 16% to 17% for non-natives, which is considerably higher than the average take-up rates observed in Table 3.

Table 5 presents results that combine stated private non-group coverage with public coverage to account for the aforementioned potential misclassification problem and private group insurance. Private non-group alone 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, as noted 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 non-natives there are decreases in private group coverage and some evidence of higher public/non-group coverage associated with the SCHIP expansions; the results suggest that when accounting for potential reporting errors by non-natives public coverage might be 2-3 percentage points higher. By contrast for natives the differences between the effect of SCHIP on public coverage in Tables 3 and 5 are slight, suggesting little evidence of misreporting. In addition, private health insurance substitution might be more of an issue for non-natives relative to natives. Waiting period effects are generally consistent with our expectations.

Implied crowd-out estimates for non-natives range from 7% to 38% without correcting for possible misclassification of non-group coverage and 19% to 57% when correcting estimates for potential misclassification. For natives the estimates range from 13% to 26% without correction and 3% to 52% with correction.

**Discussion**

While the public insurance expansions of the past two decades have been generally successful, a high fraction of eligible children remain uninsured. Devising strategies to increase enrollment among eligible children requires an understanding of how take-up varies within the population. Children of immigrants are a vulnerable group who face special barriers and challenges with respect to public insurance enrollment. Previous research suggests that the Medicaid expansions of the late 1980s and early 1990s did little to increase insurance coverage or access to care for this group. In this paper we test for differential impacts of eligibility for the State Children's Health Insurance Program on children of native-born and foreign-born adults. Contrary to the prior research, we find significant program take-up for both groups. 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.

A key feature of most SCHIP programs is the requirement that children be uninsured for some period of time before enrolling in the program. For children of both native and non-native adults, we find that these waiting periods are effective at reducing the substitution of public insurance for private coverage and, by inhibiting such transitions, limit the number of children gaining public coverage. Crowd-out estimates were generally similar between the two groups, though they were somewhat sensitive to the modeling assumptions.

Earlier work on SCHIP indicated that there is some misclassification of insurance coverage in survey data, with some respondents whose children are enrolled in SCHIP plans reporting non-group private coverage. Our results suggest that this classification error is more of an issue for non-native families. This is perhaps not surprising given that non-native families are likely to have less familiarity with state programs and the US health insurance system overall.

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, CPS Data 1997-2001**

| | 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**

| | | | | | | Change |
|---|---|---|---|---|---|---|
| | | | ***A. All Children*** | | | |
| **Native** | **1996** | **1997** | **1998** | **1999** | **2000** | **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** | | | | | | |
| Public Eligibility | 0.436 | 0.427 | 0.605 | 0.630 | 0.580 | 0.144 |
| Public Coverage | 0.269 | 0.248 | 0.255 | 0.263 | 0.280 | 0.011 |
| Private Coverage | 0.470 | 0.487 | 0.482 | 0.488 | 0.513 | 0.042 |
| Uninsured | 0.293 | 0.292 | 0.288 | 0.269 | 0.241 | -0.052 |
| | | | ***B. Family Income less than FPL*** | | | |
| | | | | | | **Change** |
| **Native** | **1996** | **1997** | **1998** | **1999** | **2000** | **1996 to 2000** |
| Public Eligibility | 0.943 | 0.967 | 0.998 | 1.000 | 1.000 | 0.057 |
| Public Coverage | 0.616 | 0.580 | 0.549 | 0.561 | 0.580 | -0.036 |
| Private Coverage | 0.260 | 0.265 | 0.301 | 0.309 | 0.309 | 0.050 |
| Uninsured | 0.182 | 0.203 | 0.216 | 0.197 | 0.171 | -0.011 |
| **Non-Native** | | | | | | |
| Public Eligibility | 0.939 | 0.952 | 0.999 | 1.000 | 1.000 | 0.061 |
| Public Coverage | 0.497 | 0.492 | 0.458 | 0.482 | 0.481 | -0.016 |
| Private Coverage | 0.171 | 0.191 | 0.194 | 0.195 | 0.200 | 0.029 |
| Uninsured | 0.365 | 0.352 | 0.368 | 0.345 | 0.346 | -0.019 |
| | | | ***C. Family Income 100% to 300% of FPL*** | | | |
| | | | | | | **Change** |
| **Native** | **1996** | **1997** | **1998** | **1999** | **2000** | **1996 to 2000** |
| Public Eligibility | 0.135 | 0.149 | 0.333 | 0.460 | 0.480 | 0.346 |
| Public Coverage | 0.126 | 0.114 | 0.116 | 0.143 | 0.159 | 0.033 |
| Private Coverage | 0.776 | 0.773 | 0.777 | 0.766 | 0.766 | -0.009 |
| Uninsured | 0.124 | 0.134 | 0.129 | 0.117 | 0.102 | -0.022 |
| **Non-Native** | | | | | | |
| Public Eligibility | 0.138 | 0.153 | 0.550 | 0.628 | 0.574 | 0.436 |
| Public Coverage | 0.150 | 0.134 | 0.179 | 0.203 | 0.252 | 0.102 |
| Private Coverage | 0.570 | 0.571 | 0.541 | 0.538 | 0.551 | -0.019 |
| Uninsured | 0.317 | 0.325 | 0.311 | 0.281 | 0.242 | -0.075 |

**Table 3: 2SLS Estimates of the Effect of Public Insurance Eligibility on Insurance Coverage for Children 300% FPL or Below**

| | Natives | Non-Natives | Natives | Non-Natives |
|---|---|---|---|---|
| **A. Public Insurance** | | | | |
| Eligible for Public Insurance | 0.067*** | 0.103*** | 0.061* | 0.123*** |
| | (0.023) | (0.037) | (0.034) | (0.037) |
| $R^2$ | 0.283 | 0.205 | 0.287 | 0.222 |
| **B. Private Insurance** | | | | |
| Eligible for Public Insurance | –0.011 | –0.039 | –0.015 | –0.025 |
| | (0.022) | (0.033) | (0.033) | (0.040) |
| $R^2$ | 0.259 | 0.217 | 0.264 | 0.231 |
| **C. Uninsurance** | | | | |
| Eligible for Public Insurance | –0.058** | –0.088* | –0.045 | –0.114** |
| | (0.019) | (0.047) | (0.038) | (0.052) |
| $R^2$ | 0.040 | 0.082 | 0.050 | 0.096 |
| **State x Year Interaction** | N | N | Y | Y |
| **Number of Observations** | 85,946 | 23,194 | 85,946 | 23,194 |

Notes: Results are based on regressions using March CPS data from 1997 to 2001. All regressions control for pubic eligibility (instrumented with simulated eligibility), state AFDC/TANF caseload, state unemployment rate, GSP, number of persons in the family, whether a single mother or father is present in the household versus two parent household, gender, race (black, Hispanic, other) interacted with time indicators, number of workers in the household, whether anyone in the household works for a large firm indicators for the number of people in the household with some college education interacted with time indicators, number of people in the house in fair or poor health, residence in an urban area, child age indicators, and state and time indicators. Standard errors clustered by state are in parentheses.

* $0.05 < p < 0.10$

** $0.01 < p < 0.05$

*** $p < 0.01$

**Table 4: 2SLS Estimates of the Effect of Public Insurance Eligibility and Probationary Periods on Insurance Coverage for Children 300% FPL or Below**

| | Natives | Non-Natives | Natives | Non-Natives |
|---|---|---|---|---|
| **Public Insurance** | | | | |
| Eligible for Public Insurance | 0.131*** | 0.165*** | 0.087** | 0.160*** |
| | (0.032) | (0.049) | (0.039) | (0.043) |
| Waiting Period (months) | –0.021*** | –0.018*** | –0.019*** | –0.018*** |
| | (0.004) | (0.006) | (0.004) | (0.006) |
| $R^2$ | 0.297 | 0.211 | 0.294 | 0.226 |
| **Private Insurance** | | | | |
| Eligible for Public Insurance | –0.051* | –0.103* | –0.032 | –0.057 |
| | (0.029) | (0.061) | (0.037) | (0.059) |
| Waiting Period (months) | 0.013*** | 0.019 | 0.013*** | 0.016 |
| | (0.004) | (0.011) | (0.004) | (0.011) |
| $R^2$ | 0.271 | 0.232 | 0.229 | 0.239 |
| **Uninsurance** | | | | |
| Eligible for Public Insurance | –0.079** | –0.089 | –0.052 | –0.118* |
| | (0.032) | (0.064) | (0.043) | (0.063) |
| Waiting Period (months) | 0.007** | 0.0002 | 0.005 | 0.002 |
| | (0.003) | (0.007) | (0.004) | (0.008) |
| $R^2$ | 0.034 | 0.082 | 0.048 | 0.095 |
| **State x Year Interaction** | N | N | Y | Y |
| **Number of Observations** | 85,946 | 23,194 | 85,946 | 23,194 |

Notes: Results are based on regressions using March CPS data from 1997 to 2001. All regressions control for pubic eligibility and waiting periods (instrumented with simulated eligibility and simulated eligibility*waiting period), as well as state AFDC/TANF caseload, state unemployment rate, GSP, number of persons in the family, whether a single mother or father is present in the household versus two parent household, gender, race (black, Hispanic, other) interacted with time indicators, number of workers in the household, whether anyone in the household works for a large firm indicators for the number of people in the household with some college education interacted with time indicators, number of people in the house in fair or poor health, residence in an urban area, child age indicators, and state and time indicators. Standard errors clustered by state are in parentheses.

* $0.05 < p < 0.10$

** $0.01 < p < 0.05$

*** $p < 0.01$

**Appendix Table A-1. SCHIP Expansions by State**

| State | Date Implemented | %FPL Eligibility Limit, 1-5 yrs | | %FPL Eligibility Limit, 15 yrs | |
|---|---|---|---|---|---|
| | | 1996 | 2000 | 1996 | 2000 |
| AK | March 1999 | 133% | 200% | 71% | 200% |
| AL | February 1998 | 133% | 200% | 15% | 200% |
| AR | October 1998 | 133% | 200% | 18% | 200% |
| AZ | October 1997 | 133% | 200% | 30% | 200% |
| CA | July 1998 | 133% | 200% | 82% | 200% |
| CO | April 1998 | 133% | 185% | 37% | 185% |
| CT | October 1997 | 185% | 300% | 81% | 300% |
| DC | September 1997 | 133% | 200% | 50% | 200% |
| DE | October 1998 | 133% | 200% | 100% | 200% |
| FL | April 1998 | 133% | 200% | 28% | 200% |
| GA | September 1998 | 133% | 200% | 100% | 200% |
| HI | January 2000 | 133% | 185% | 100% | 100% |
| IA | July 1998 | 133% | 185% | 37% | 185% |
| ID | October 1997 | 133% | 150% | 29% | 150% |
| IL | January 1998 | 133% | 185% | 46% | 185% |
| IN | October 1997 | 133% | 150% | 27% | 150% |
| KS | July 1998 | 133% | 200% | 100% | 200% |
| KY | July 1998 | 133% | 200% | 33% | 200% |
| LA | November 1998 | 133% | 150% | 10% | 150% |
| MA | October 1997 | 133% | 200% | 86% | 185% |
| MD | July 1998 | 185% | 200% | 40% | 200% |
| ME | July 1998 | 133% | 185% | 125% | 200% |
| MI | May 1998 | 150% | 200% | 150% | 200% |
| MN | September 1998 | 275% | 280% | 275% | 280% |
| MO | October 1997 | 133% | 300% | 100% | 300% |
| MS | July 1998 | 133% | 200% | 34% | 200% |
| MT | January 1998 | 133% | 150% | 41% | 150% |
| NC | October 1998 | 133% | 200% | 100% | 200% |
| ND | October 1998 | 133% | 140% | 40% | 140% |
| NE | May 1998 | 133% | 185% | 33% | 185% |
| NH | May 1998 | 185% | 300% | 185% | 300% |
| NJ | February 1998 | 133% | 350% | 41% | 350% |
| NM | March 1999 | 185% | 235% | 185% | 235% |
| NV | October 1998 | 133% | 200% | 31% | 200% |
| NY | April 1998 | 133% | 192% | 51% | 192% |
| OH | January 1998 | 133% | 150% | 33% | 150% |
| OK | December 1997 | 133% | 185% | 100% | 185% |
| OR | July 1998 | 133% | 170% | 100% | 170% |
| PA | June 1998 | 133% | 200% | 41% | 200% |
| RI | October 1997 | 250% | 250% | 51% | 250% |
| SC | October 1997 | 133% | 150% | 48% | 150% |
| SD | July 1998 | 133% | 140% | 100% | 140% |
| TN | October 1997 | 400% | 400% | 100% | 400% |
| TX | July 1998 | 133% | 133% | 17% | 100% |
| UT | August 1998 | 133% | 200% | 100% | 200% |
| VA | October 1998 | 133% | 185% | 100% | 185% |
| VT | October 1998 | 225% | 300% | 225% | 300% |
| WA | January 2000 | 133% | 250% | 200% | 250% |
| WI | April 1999 | 133% | 185% | 45% | 185% |
| WV | July 1998 | 133% | 150% | 100% | 150% |
| WY | April 1999 | 133% | 133% | 55% | 133% |