Economic Research Initiative on the Uninsured
Working Paper Series

DECOMPOSING THE GROWING DISPARITY IN HEALTH INSURANCE COVERAGE BETWEEN HISPANICS AND NON-HISPANIC WHITES

Matthew S. Rutledge
Catherine G. McLaughlin
University of Michigan

ERIU Working Paper 56
http://www.umich.edu/~eriu/pdf/wp56.pdf

Economic Research Initiative on the Uninsured
University of Michigan
109 S. Observatory, M3242, SPH-II
Ann Arbor, MI 48109-2029

March, 2008

Abstract

The percentage of non-Hispanic Whites without health insurance has fallen slightly since the early 1980's, and the rate for other racial groups has remained stable. Meanwhile, the uninsured rate among Hispanics has increased dramatically. In this paper, we estimate how differences between the two groups in citizenship, education, public coverage eligibility, income, and labor market participation have contributed to the increasing gap in health insurance coverage. Pooling panels of the Survey of Income and Program Participation to gather over twenty years of monthly data, we estimate the extent of the contribution of these factors using two different methods: (1) the statistical significance of the trend line in the uninsured rate gap, regression-adjusted for each variable sequentially, and (2) pairwise Blinder-Oaxaca-style decomposition of the change in the gap between 1984 and each year thereafter. We find that, consistent across decomposition method, more than 70 percent of the increase in the gap, or between 2.5 million and 4.9 million extra uninsured Hispanics, remains unexplained. While differences in observables account for most of the divergence in the public coverage and employer-sponsored coverage rates, a substantial portion of the relative decline in coverage through a family member remains unexplained. We suggest that job composition, social networks, risk tolerance, and healthy quality may, even controlling for factors that the literature has shown to explain coverage disparities, make Hispanics (particularly those of Mexican origin) less likely to seek, or be able to find, health insurance coverage than white non-Hispanics.

1 Introduction

Between 1987 and 2003, the number of Americans without health insurance grew from 31.0 million to 43.4 million, or from 12.9 percent of the population to 15.1 percent (DeNavas-Walt et al, 2007).[1] This expansion in the ranks of the uninsured, despite public coverage expansions and a macroeconomy that was generally booming during this time, is a concern for policymakers, as lower coverage rates could result in more exposure to medical and financial risk and less access to necessary health care. The growing uninsured rate has therefore received a great deal of attention both in the mainstream media and by academic journals in the fields of medicine, public health, and economics.

Less recognized has been the fact that for many population subgroups, the uninsured rate has remained constant, or actually declined, over this time. In 1983, 19 percent of children under age 18 were uninsured; in 2003, this rate had fallen to fifteen percent, aided by expansions to Medicaid and SCHIP.[2] The number of elderly (over 65) uninsured remains negligible, due to pervasive Medicare coverage. The uninsured rate fell even among some groups of nonelderly adults, who are largely ineligible for public coverage. Though the uninsured rate for Blacks remains ten percentage points higher than Whites, the proportion of Blacks without health insurance fell slightly, mirroring the small decline in uninsured Whites. Asians and Native Americans are also slightly less likely to be uninsured than they were in the 1980's.

For Hispanics, though, the health insurance problem has only grown worse. Already sixteen percentage points higher than non-Hispanic Whites in the early 1980's, the uninsured rate for Hispanics increased to 40.3 percent in 2003, a gap of nearly 27 percentage points with non-Hispanic Whites. Whereas the uninsured rates for Whites and Blacks have fluctuated, with an increase in the early 1990's sandwiched by declines during the late 1980's and late 1990's, the Hispanic uninsured rate has grown almost monotonically throughout this period. While much of the growth in the number of uninsured Hispanics has been concentrated in the non-native population, U.S.-born Hispanics are also more likely to be uninsured compared to twenty years ago.

In a previous paper (Rutledge and McLaughlin, 2008), we analyze the trends in the uninsured rate across racial and ethnic groups, and within the Hispanic group, in more detail. We find that the overall uninsured rate would have been essentially flat over the last two decades if the Hispanic noncitizen proportion of the population had remained constant, and actually would have declined if the proportion Hispanic (without regard to citizenship) had remained unchanged. We also suggest that changes in the composition of the Hispanic population, particularly the growth in the likelihood of having less than a high school education relative to other ethnic groups, have contributed to the growing gap

[1] These figures are from the Current Population Survey.

[2] These figures are from the Survey of Income and Program Participation.

between Hispanics and non-Hispanics.

This paper extends our previous work on uninsured rate trends by formally analyzing how differences in citizenship status, educational attainment, public coverage eligibility, income, and labor market outcomes between Hispanics and non-Hispanic Whites have contributed to the divergence in uninsured rates between the two groups. We examine not only how observable differences contribute to point-in-time coverage gaps, but also how these observables have contributed to the *growth* in the gap over two decades.

We use two methods to decompose the contribution of observable variables to the diverging uninsured rates: 1) estimation of a regression-adjusted trend line, from the coefficients of the interaction of the Hispanic indicator with year dummies, and determining how this trend line changes when adding each variable sequentially, and 2) Blinder-Oaxaca-style decomposition of the trend from a base period to the end of our sample. We find that, consistent across our decomposition methods, differences in citizenship and educational attainment between Hispanics and non-Hispanic Whites account for each about one quarter of the growth in the uninsured rate gap, and a similar amount of the annual point-in-time gaps, more than any other of the characteristics we observe. Still, in each of our decomposition methods, more than 70 percent of the divergence in uninsured rates from 1984 to 2003 remains unexplained.

Though the growth in the uninsured rate among Hispanics was steady throughout our sample period, the degree to which we can explain the growing gap differs between the first and second halves of our sample window. In our analysis, we find that observables can account for much of the uninsured rate divergence between 1984 and 1995; the growth in the gap is reduced by between 40 and 75 percent during this time, mostly due to Hispanic-White differences in citizenship. Between 1996 and 2003, however, our regression-adjusted gap is actually larger than the actual gap; that is, Hispanic-White differences in citizenship and public eligibility actually suggest that the gap should be larger than it is during this time.

We also analyze the trends in the rate of coverage separately by insurance source. The gap in public coverage has been reliably countercyclical, but after adjusting for observables, Hispanic and white non-Hispanic public coverage rates are essentially equal and constant over time. Adjusting for observables, Hispanics were actually *more* likely to have private coverage through their own employer during the early years of our sample, but were increasingly less likely to obtain employer sponsored coverage in late 1990's and early 2000's. There was also a sharp decline for Hispanics in coverage through another family member, the one insurance source category where a significant portion of the gap with non-Hispanic Whites remains unexplained even after controlling for observable characteristics.

Realizing that the Hispanic population is far from homogeneous, we repeat the decomposition exercises for subgroups by country of origin. We find that the uninsured rate gap between those of Mexican heritage and non-Hispanic Whites is significantly larger than for Puerto Ricans and Cubans, and that most of the growth in this gap remains unexplained. In addition, we are also unable to account for the divergence between Mexicans and these other two Hispanic subgroups, controlling for intra-Hispanic differences in the variables in our data.

In 2003, 9.9 million Hispanics between the ages of 18 and 55 were uninsured, according to our sample. If the uninsured rate gap had remained constant at its 1984 level, 6.4 million nonelderly Hispanic adults would have been uninsured. Based on our estimations, only 0.9 million of this 3.4 million difference can be explained by observable characteristics, leaving an unaccounted-for increase of 2.5 million uninsured Hispanic adults since 1984. Similarly, if the proportion of the population that was Hispanic had remained constant at 1984 levels but the uninsured gap was allowed to grow as it actually did, 4.9 million extra Hispanic adults remain unexplained in our estimations.

Finally, as much of the increase in the uninsured rate for Hispanics relative to non-Hispanic Whites remains unexplained, we offer some potential reasons for the widening gap, among them job composition, social networks, risk tolerance, and health quality.

In section 2, we relate our work to previous literature, which documents the growth in the uninsured rate and points out variables that have had the most impact on this growth. We describe our data in section 3, and motivate our decompositions by displaying trends in insurance coverage over our sample window in section 4. In section 5, we describe the two decomposition methods, and outline results in section 6. We discuss the magnitude of our findings and suggest additional explanations for the growth of the uninsured rate gap in section 7. Section 8 concludes.

2 Previous Literature

The concentration of the growth in the uninsured within the Hispanic population is not a new finding. Other researchers have attributed much of the increase in the uninsured rate to the growing Hispanic population and their increased likelihood to be without health insurance coverage. Carrasquillo, Himmelstein, et al (1999) analyzed trends in the uninsured rate from 1989 to 1996 and found that Hispanics account for 36.4 percent of the increase in the number of uninsured during that time. Rutledge and McLaughlin (2008) calculate that the growth in the Hispanic share of the population accounts for 31 percent of the increase in the number of uninsured adults.

Though we focus on how the gap in uninsured rates between Hispanics and non-Hispanics has changed over time, we begin by decomposing how personal characteristics contribute to the gap within each year. Our choice of variables that we think will be important owes a great deal to Monheit and Vistnes (2000), who do exactly this point-in-time decomposition. They use Blinder-Oaxaca decompositions to find that most of the gap between Hispanic and white men can be explained by wages, family income, and education, though much of the gap remains unexplained.

Two other studies suggest separate analysis and decomposition by insurance source. Waidmann, Garrett, and Hadley (2000) decompose how demography, human capital, labor and health care market changes, and state policies affect offers and takeup of employer sponsored insurance (ESI), and find that most of the gap is explained by English language proficiency and education. Buchmueller, LoSasso, Lurie, and Dolfin (2007) sequentially add human capital and employer variables (similar to our Method 1) to regressions of ESI eligibility, offers, and takeup, and find that only the gap between natives and non-citizens in offer rates remains substantially unexplained.

Other studies make clear that the Hispanic population is not uniform, suggesting separate analysis by country of origin. In descriptive work, Shah and Carrasquillo (2006) examine trends over a twelve-year period among Latino subgroups compared to non-Hispanic whites, and find that Mexicans were hardest hit by Medicaid cuts in the 1990's and employer coverage declines in the early 2000's. Berk, Albers, and Schur (1996) find similar results between 1977 and 1992. Fronstin, Goldberg and Robins (1987) use Blinder-Oaxaca decomposition to compare Mexicans to Puerto Ricans and Cubans, both of whom have lower uninsured rates, and find that most of the difference can be explained by wages, age, education, industry, and firm size.

Besides county of origin, Hispanics differ in other ways, including nativity, gender, and employment, that influence their coverage probability. Hamilton, Hummer, You, and Padilla (2006) separate Mexican-Americans by generation; while second-generation Mexican-Americans have much lower coverage rates than blacks, third-generation Mexican-Americans actually have higher rates than blacks and are similar to whites. Alegria et al. (2005) find gender differences for Latinos, with females far more likely to have public coverage and males more likely to have private coverage, and that only immigrants within their first five years in the U.S. have substantially different coverage rates than U.S. natives. Schur and Feldman (2000) look at firm size and industry, and find that even within a finely-defined industry, Hispanics have lower coverage rates. While we are unable to directly control for generation, we do include gender, taking into account that public insurance programs may affect women differently from men, and industry.

Finally, it is impossible to discuss public policy involving Hispanics without addressing immigration. Camarota and Edwards (2000) attribute much of the growth in the overall uninsured rate to immigrants, due mostly to lower educational attainment and higher poverty rates. Borjas (2003) found no significant change in the uninsured rate among non-citizens following welfare reform in the mid-1990's, as reductions in public coverage eligibility were cancelled out by increased labor force participation and a greater rate of employer-sponsored coverage. Kandilov (2007), though, finds no difference in private coverage rates, and higher uninsured rates overall, among immigrants in states with more restrictive Medicaid rules post-reform. Our study takes into account not only how citizenship impacts the likelihood of insurance coverage, but how citizenship can differentially affect Hispanics and non-Hispanics.

Our estimates include variables that these studies have found to be important in similar decomposition methods. This paper contributes to the discussion by analyzing the growth in the uninsured rate gap between Hispanics and non-Hispanics over a longer period of time. We use two different decomposition methods, including Blinder-Oaxaca decomposition of the growth in the uninsured rate gap between two different points in time, a method heretofore underutilized in the data.[3] These decompositions consider multiple start and end points and take into account each year in between; earlier papers (such as Monheit and Vistnes 2000) that analyzed changes in the uninsured rate were forced to analyze fixed start and end years corresponding to their cross-sectional samples, without regard to intermittent points in time. We perform separate trend decompositions by insurance source. We also extend our trend analysis to coverage rate differences between Hispanic subgroups, an extension and update of Fronstin, Goldberg, and Robins (1987).

3 Data

We use data from the Survey of Income and Program Participation (SIPP), which interviews households about demographics, income sources, welfare, household and family structure, jobs and work history, and health and health insurance. Interviews are conducted by survey takers every four months about each individual in the household for each separate month between interviews, so we have data for person-months. Each year from 1984 through 1993, and then again in 1996, 2001, and 2004, the Census Bureau began a new panel. Questions are comparable across panels, though over time some answer options changed (such as educational attainment). The samples sizes and length of the panels vary (Table 1). Most of the panels track respondents for close to three years, or about 8 interview waves, though the 1996 panel lasted for four years. The 1996 panel was also the largest sample, while the 1984 sample was the smallest.

[3]Three exceptions are Van Hook, Brown and Kwanda (2004) for trends in poverty among the children of immigrants, Fairlie and Sundstrom (1999) for the racial unemployment gap, and Ashraf (1996) for the gender wage gap.

Table 1: Survey of Income and Program Participation (SIPP) Panels

Panel	Waves	Sample Size	Beginning Month	Ending Month
1984	9	32,555	June 1983	July 1986
1985	8	22,088	October 1984	July 1987
1986	7	18,668	October 1985	March 1988
1987	7	18,701	October 1986	April 1989
1988	6	9,325	October 1987	December 1989
1990	8	25,641	October 1989	August 1992
1991	8	10,513	October 1990	August 1993
1992	9	21,727	October 1991	December 1994
1993	9	13,790	October 1992	December 1995
1996	12	57,096	December 1995	February 2000
2001	9	40,408	October 2000	December 2003

Note: Sample size reflects number of nonelderly adults in the first month of Wave 1 of that panel.

Our dataset pools each completed panel[4] into one large sample that is continuous for all but seven months (March through September 2000, between the end of the 1996 panel and the beginning of the 2001 panel) of the interval between June 1983 and December 2003.

We focus on individuals age 18 to 64, since the difference in the uninsured rate between Hispanics and white non-Hispanics is most pronounced for this age group. The uninsured rate among children of both groups has been stable over this time (Appendix Figure 1). Because of pervasive Medicare coverage, the uninsured rate for the elderly is negligible throughout, though Hispanics are more likely to be without health insurance among the elderly as well (Appendix Figure 2).

In our regressions, we exclude non-Hispanic blacks, Asians, and Native Americans from the sample. Our goal is to explain the difference in the trend in the uninsured rate between Hispanics and the majority group, white non-Hispanics, where the gap is most stark. We obtain no additional explanatory power by the inclusion of individuals who are in these other groups.[5] We still include race indicators in the regressions to account for potential differences between white and black Hispanics (or Asian Hispanics, or Native American Hispanics, though these interactions with Hispanic are much less common), though the results without these variables are not substantially different.

Our primary dependent variable is an indicator variable for whether the individual reported health insurance coverage, irrespective of source, in the sur-

[4]The 1989 panel was discontinued after three waves, and we chose not to include it in our sample. Also, as the 2004 panel is still in progress and subject to revision and imputation, we chose not to include it either.

[5]The coefficients and standard errors in regressions including all blacks, Asians, and Native Americans are nearly identical.

veyed month.[6] We consider individuals with missing values for the insurance questions to be out-of-sample; in our time-series figures, we correct the SIPP-provided person-month weights to account for potential non-random attrition (see Appendix), though all regressions are unweighted.[7]

As we will explain in more detail in section 4, in this paper we choose to represent the individual's insurance coverage choice as a linear model with only this one outcome, insured or uninsured. We could (and have) run multivariate regressions where the possible coverage outcomes are private, public, or none. The problem with this structure is that most individuals in our sample are not eligible for public coverage, since Medicaid is typically available only to expectant mothers, the elderly, and the disabled, so public coverage is an appropriate part of the choice set of only a select few. Unfortunately, we do not know with certainty who is eligible for public coverage. One option would be to separate those who are eligible and run a multivariate regression for them, while leaving the remainder in a more traditional linear or nonlinear univariate choice model. This would be our choice of strategies if we had a clear indicator of public coverage eligibility. Even if we knew the eligibility rules in the individual's state, we would likely have a large number of individuals who appear to be ineligible but have public coverage.[8]

A simpler option, and the one we ultimately choose, is to control for public coverage eligibility in the regression where uninsured is the dependent variable. Our proxy for public coverage eligibility is an indicator variable equal to unity when the individual is below the (time-varying) federal poverty line. We also interact this variable separately with our female and Hispanic indicators; female because poor women are more likely to be eligible than poor men, Hispanic because we hypothesize that Hispanics may have different takeup responses to public coverage eligibility than non-Hispanics. Also, we further narrow down

[6]SIPP asks about insurance status for each month individually; while there could be seam bias, where the retrospective months are less reliable than the interview month, our results are no different using only the interview months. The Current Population Survey (CPS) uninsured rates are more often quoted in the press, but the CPS asks whether the individual ever had health insurance during the previous year; this is a very different question than the SIPP, as insurance status often changes during the calendar year. Researchers have found that the CPS measure of insurance status actually follows the point-in-time status around the interview month more closely than the ever-in-year status implied by the question. We repeat our analysis for the CPS and get similar results, though CPS did not ask about citizenship status until 1993.

[7]The coefficients in weighted regressions, using either the uncorrected or corrected weights, are nearly identical to the unweighted regressions.

[8]The other kind of error, where individuals who appear to be eligible do not have public coverage, are considered "conditionally covered" in the health insurance literature. While hospitals will assist them in acquiring their rightful public assistance when they present for emergency care, they miss out on regular preventive and non-emergency care. Because of this difference in what kind of care is covered, there is a debate over whether these individuals should be considered insured or uninsured. The definition of "covered by health insurance" that is consistent with the SIPP questionnaire would include non-emergency care, so we consider the "conditionally covered" to be uninsured.

our sample to just those who are between the ages of 18 and 55, since a nontrivial number of respondents between the ages of 56 and 64 report Medicare coverage.[9]

In addition, SIPP provides data on the source of insurance. Though the exact options on the questionnaire change between panels, we are able to construct a consistent indicator for whether an individual has coverage through a public program (grouping Medicare, Medicaid, SCHIP, and other federal and state programs together) or through a private insurer. We further divide private coverage into coverage obtained through one's own employer, coverage obtained in one's own name from some source other than own's employer, and coverage obtained in someone else's name (irrespective of who that person is, or where that person obtains coverage). We separately analyze trends by each insurance source category, and separately decompose how the same right-hand side variables contribute to the Hispanic-White gap in each insurance source.

4 Trends in the Uninsured Rate

The uninsured rate among adults aged 18 to 55 in the U.S. increased slightly, though significantly, between 1983 and 2003 (Figure 1). The uninsured rate increased from 18 percent in the mid 1980's to almost 20 percent in 1993 and 1994, before falling during the economic boom of the late 1990's. After 1998, the proportion uninsured increased again, resulting in an overall two-decade upward trend of 0.08 percentage points per year, which is statistically significant at the 99 percent confidence level.

The overall uninsured rate increased despite significant downward trends among non-Hispanic Whites and non-Hispanic Blacks, particularly since the early 1990s. Both groups follow the pattern of the overall rate, but the larger decreases in the latter decade led to a decline in the uninsured rate for Whites of 0.1 percentage points per year, and 0.07 percentage points for Blacks. The uninsured rate among Asians (not shown) had no significant trend, while among Native Americans (also not shown), there was a statistically significant downward trend of nearly 0.4 percentage points per year.

Alone among the racial and ethnic groups investigated, the uninsured rate for Hispanics trended upward, and by amounts that dwarf the improvement in the other groups. The Hispanic uninsured rate was 27.8 percent in 1984, more than twelve percentage points greater than the uninsured rate for non-Hispanic

[9] While the percent of public coverage recipient above 100 percent of the federal poverty line increased slightly during our sample window, the vast majority of public coverage recipients had family incomes below the poverty line. Throughout our sample period, the proportion of the population under the poverty line was greater than the proportion with public coverage, indicating a large number of "conditionally covered" (see note 8). Also, the percent uninsured closely matches the difference between our eligibility proxy and the percent with public coverage, suggesting that the "conditionally covered" did not have any health insurance.

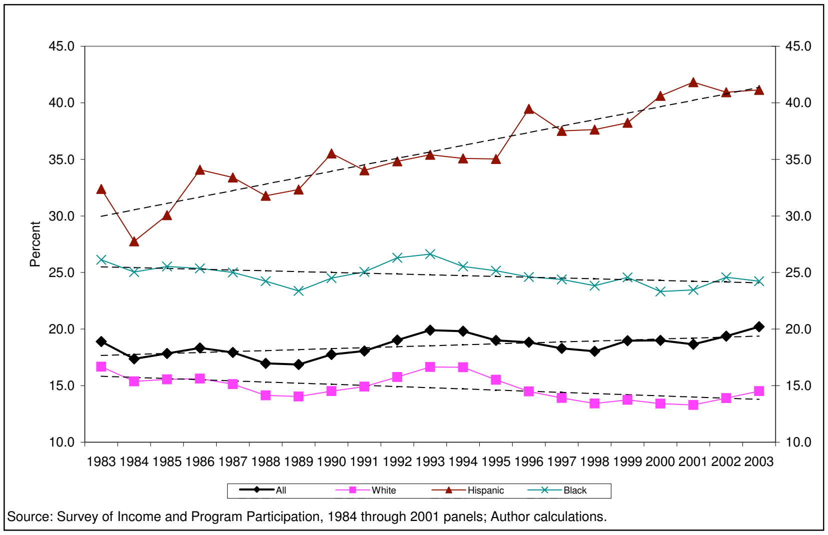


Figure 1: Uninsured Rate by Racial and Ethnic Group, Age 18 to 55

Whites. Over the next twenty years, the proportion of Hispanics without health insurance increased by 0.6 percentage points per year, to more than 40 percent in each year of the 2000's. The gap with non-Hispanic Whites more than doubled, increasing to 25 percentage points in 2003. At the beginning of our sample, a nonelderly Hispanic adult was nearly twice as likely to be uninsured as her non-Hispanic White peer; by the end of our sample, she is almost three times as likely.

Most, but not all, of the growth in the Hispanic uninsured rate was among non-citizens (Figure 2). In 1984, the uninsured rate among Hispanic non-citizens was 37.8 percent. By 2003, their uninsured rate was more than 57 percent; since 1996, non-citizens of Hispanic origin are more likely to be uninsured than insured.

While the level of insurance coverage is higher and the rate of growth smaller than for those born elsewhere, native-born Hispanics were also seeing an increase in the likelihood of being uninsured. The uninsured rate for native-born Hispanics increased by a statistically significant 0.3 percentage points per year over this time, from 24.3 percent in 1983 to 31.0 percent in 2003. The insurance coverage gap for Hispanics, then, cannot be written off as just an immigration problem.

Still, over this period, the health insurance status of all immigrants got worse. Non-Hispanic non-citizens also saw an increase in their uninsured rate,

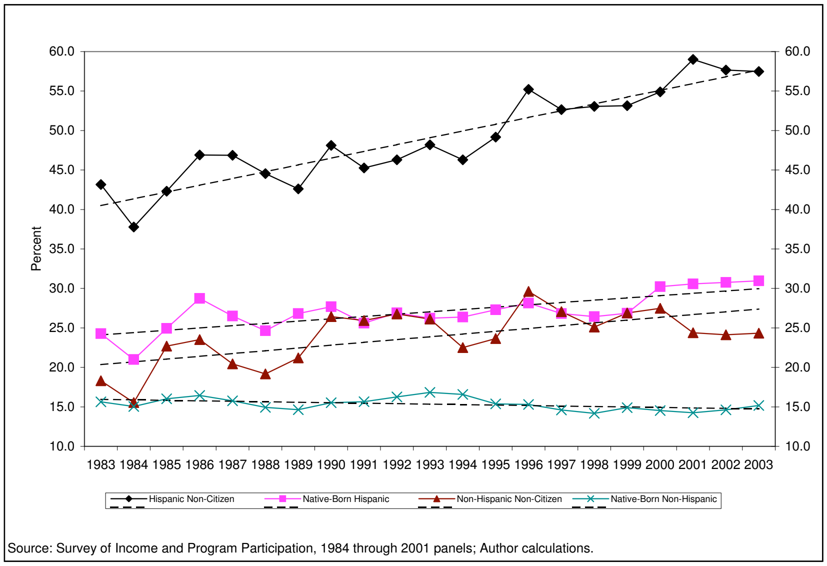


Figure 2: Uninsured Rate by Hispanic Origin and Citizenship, Age 18 to 55

from 15.5 percent in 1984 to 24.3 percent in 2003, a statistically significant upward trend of 0.4 percentage points per year; after 1996, though, the uninsured rate declines. Naturalized Hispanics (not shown) also saw a slight but significant increase, from 24.5 percent in 1983 to 28.1 percent in 2003, a slower rate of increase than Hispanic natives, while naturalized non-Hispanics had a slower, but also significant, increase.[10]

Besides citizenship, previous studies have indicated that education, income, and age explain significant portions of both the growth rate of the uninsured population and the Hispanic-non-Hispanic coverage gap.

Figures 3a (Hispanics) and 3b (non-Hispanics) display the sharp increase in the uninsured rate among those with less than a high school diploma. This increase is especially large for Hispanics, nearly a full percentage point per year, but is also significant for non-Hispanics. It is among college graduates, however, where the Hispanic gap is most stark. The uninsured rate for non-Hispanic college graduates declined from 10.4 percent to 8.5 percent from 1983 to 2003, and overall, the college graduate uninsured rate fell. For Hispanic college graduates, though, the uninsured rate actually *increased*, from less than 17 percent to more than 21 percent. Also, while the uninsured rate for non-Hispanics with just a high school diploma, or with some college experience, was nearly exactly flat over this time, the Hispanic uninsured rate for these groups significantly

[10]The uninsured rate among Hispanics without immigration information increased steadily, from around 30 percent in the mid 1980's to more than 45 percent in the 2000's.

increased.

A similar story is told in Figures 4a and 4b. Even the lowest income non-Hispanics are less likely to be uninsured in 2003 than they were in 1983, but the uninsured rate for even the highest income Hispanics has grown significantly. For each category of family income, normalized to the percent of the federal poverty line, the uninsured rate among Hispanics was already higher than non-Hispanics, and that gap has steadily grown over the last two decades.

Within age groups, the Hispanic-non-Hispanic gap has also increased (Figures 5a and 5b). The non-Hispanic uninsured rate has changed only within a tight band, with the youngest group at the highest level. The Hispanic age groups, though, have each increased, though generally at a slower rate than we see in the education and income categories.

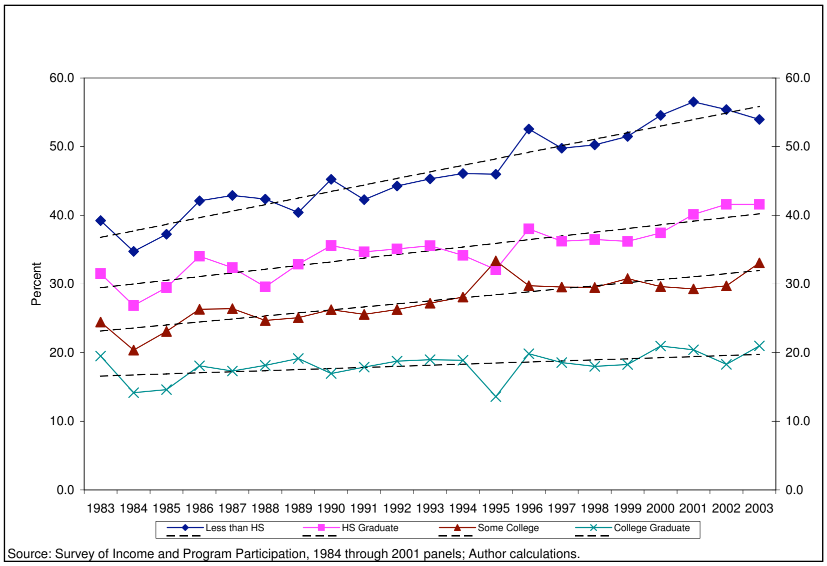
These graphs seem to indicate that differences between Hispanics and non-Hispanics in citizenship, education, and income, and to a lesser extent age,[11] have a great deal of influence on the divergence of the uninsured rate between the two groups. We will focus on these variables, along with our measure of public coverage eligibility (which is represented by the top line in Figures 4a and 4b), in our analysis.

But these variables are likely to be closely related, and their influence is hard to differentiate; obviously, greater educational attainment will generally lead to higher income, and the age-income profile is well documented. Though identifying the trends in these variables and how they relate to the uninsured rate is interesting and important, we will not be able to perceive of the relative contribution of each variable to insurance status separate from the others by merely graphing the means. Our two decomposition methods allow us to better separate the effect of each variable on the trends in the uninsured rate.

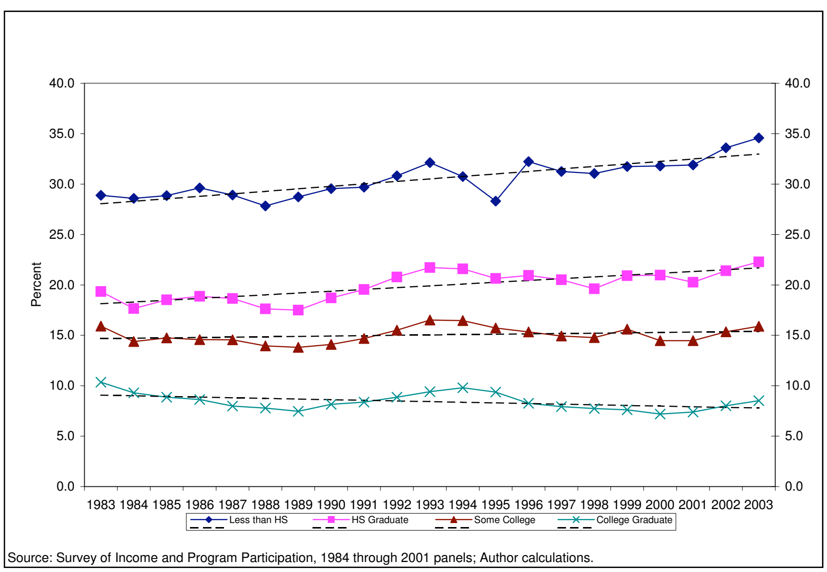
Previous studies have indicated that the Hispanic population is far from homogeneous, and that we must also consider country of origin within Hispanic ethnicity, as the uninsured rate is higher for individuals of Mexican heritage than for those from other areas in Latin America. Indeed, in each year in our sample, Mexicans had a higher uninsured rate than Puerto Ricans, Cubans, and those from other Latin countries,[12] with a higher growth rate over the two-decade period. The uninsured rate is highest for Mexican non-citizens, though Hispanic non-citizens from countries other than Mexico narrowed that gap slightly (Figure 6), particularly among Cubans. The greater growth rate for Hispanics of Mexican heritage is due to the faster increase among Mexican-Americans, while there was no significant trend among non-Mexican Hispanics born in the U.S.

[11]When the age categorical variables are added separately to Method 1, the Hispanic-year coefficients do not change much. Also, in the trend decomposition in Method 2, the effect of age is much smaller than education, citizenship, income, and public coverage. As a result, we include age only in the X variables.

[12]SIPP has consistent Hispanic heritage indicators only for Mexicans, Puerto Ricans, and Cubans, with all other Hispanics grouped together. Only in the 1996 and 2001 panels did SIPP further differentiate Hispanics.

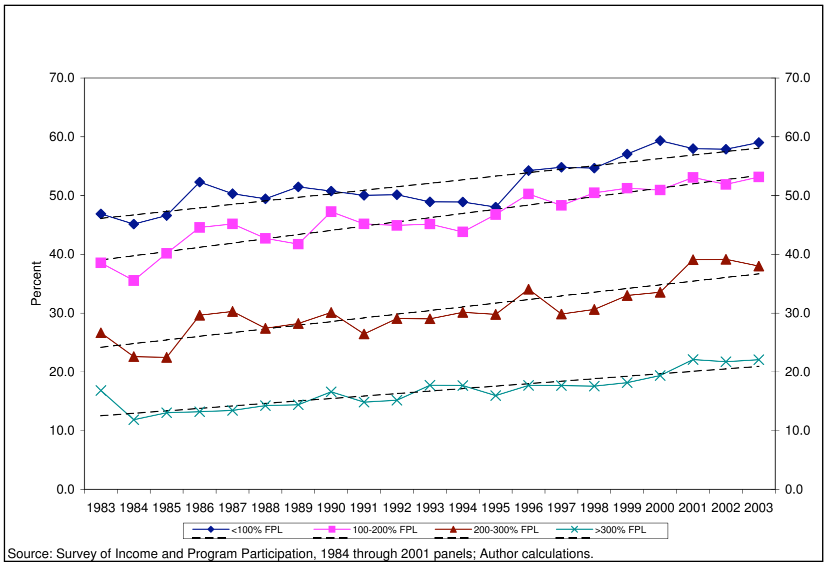


(a) Hispanics

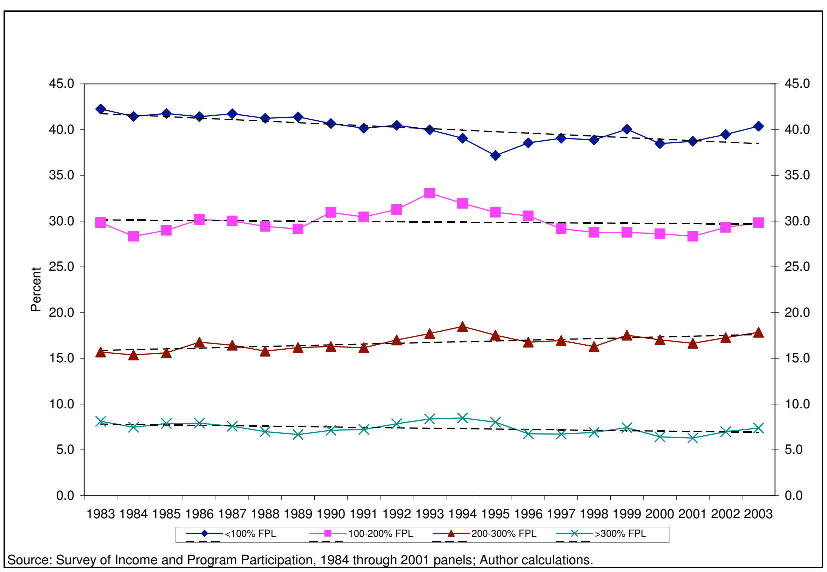


(b) Non-Hispanics

Figure 3: Uninsured Rate by Educational Attainment, Age 18 to 55

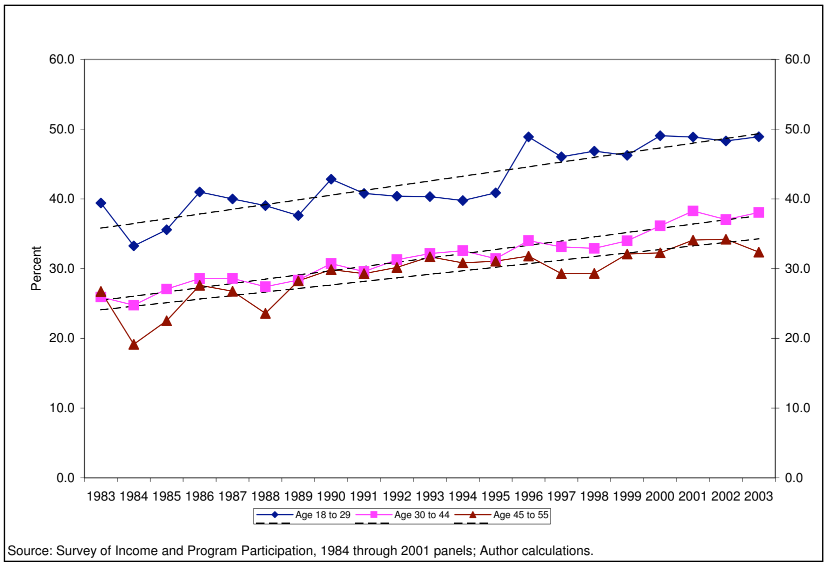


(a) Hispanics

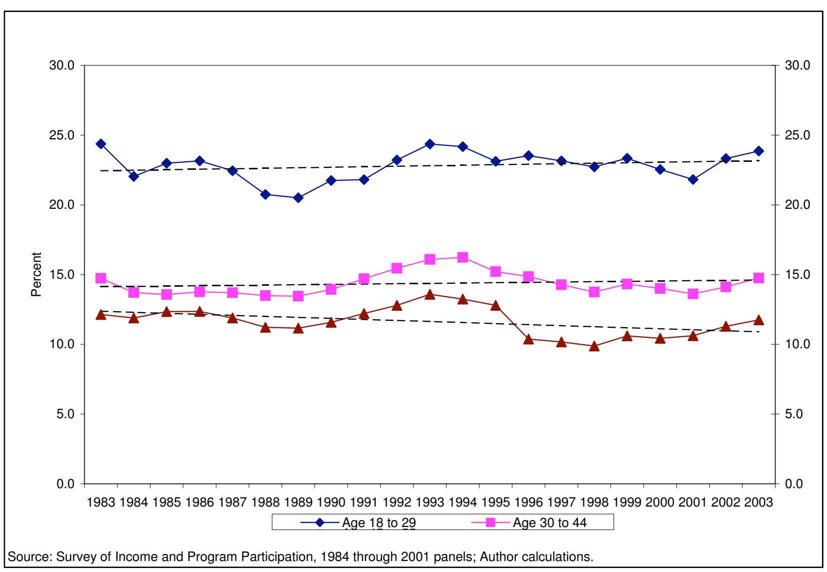


(b) Non-Hispanics

Figure 4: Uninsured Rate by Family Income as a Percent of the Federal Poverty Level, Age 18 to 55



(a) Hispanics

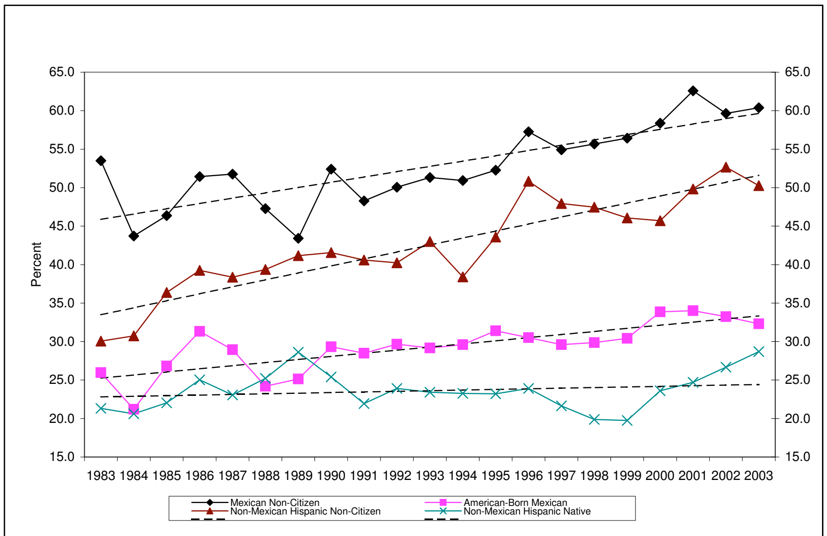


(b) Non-Hispanics

Figure 5: Uninsured Rate by Age, Age 18 to 55

Because of the very different rates of growth seen in Figure 6, we will decompose the contribution of each variable to the growth in the uninsured rate gap for Mexicans separately from other Hispanics, and also the difference between the two groups.

We may also be concerned with the source of insurance and how the coverage



Source: Survey of Income and Program Participation, 1984 through 2001 panels; Author calculations.

Figure 6: Uninsured Rate by Hispanic Subgroup and Citizenship, Age 18 to 55

type has changed over time for Hispanics relative to non-Hispanics, particularly if we feel that employer-sponsored coverage is a "better" (or at least more stable) source than public coverage or non-group private plans. From the mid-1980's through the mid-1990's, the public coverage rate increased nearly across the board, with the largest growth rate among Blacks (Figure 7). Since welfare reform in 1996, however, only Whites have maintained their public coverage level, while a much smaller percentage of Hispanics are covered by Medicare, Medicaid, or a state program. The decline was especially pronounced for Hispanic non-citizens; though the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 limited Medicaid coverage for new immigrants (Ku and Bruen, 1999), White and Black non-citizens did not see a similar drop in the public coverage rate.

Private coverage rates were relatively constant for most groups, but fell from more than 60 percent to less than 50 percent among Hispanics. There was a slight increase in the proportion with employer-sponsored insurance among Whites and Blacks, but among Hispanics, there was a substantial decrease from an already much lower percentage (Figure 8). All groups saw small (but statistically significant) and relatively equal declines in private insurance from a source other than one's employer. Finally, while the proportion of Whites receiving

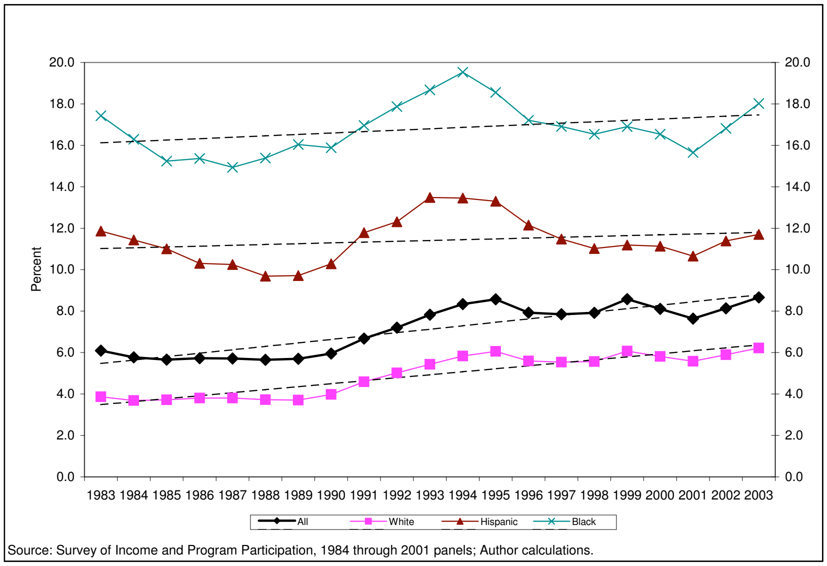


Figure 7: Public Coverage Rate By Racial and Ethnic Group, Age 18 to 55

insurance through someone else's policy declined slightly (and this percentage among Blacks actually increased, albeit insignificantly), the proportion among Hispanics fell significantly, from more than 20 percent to less than 15 percent.

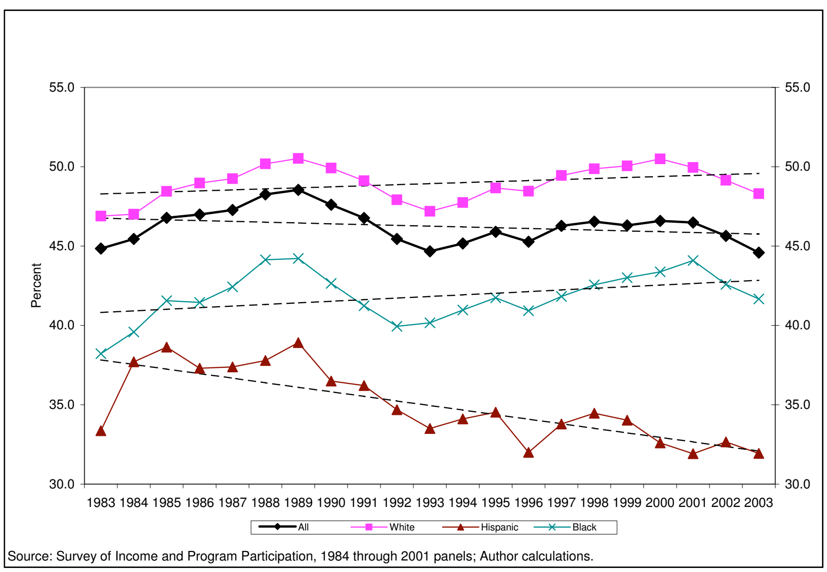


Figure 8: Employer-Sponsored Coverage in Own's Name Rate By Racial and Ethnic Group, Age 18 to 55

5 Methods

At its most basic, we run the following linear regression model, for individual i in month t:

$$Unins_{it} = \beta_0 + \beta_1 Hisp_i + \gamma' X_{it} + \sum_{y=1983}^{2002} [\zeta_y I(year_{it} = y) + \delta_y Hisp_i I(year_{it} = y)], \quad (1)$$

where $I(year_{it} = y)$ is an indicator function that equals unity when the person-month of observation occurs in year y. Thus, our "adjusted" measure of the Hispanic-White in the uninsured rate in year y, accounting for observable variables, is $\beta_1 + \delta_y$.

Our X_{it} vector in each regression includes indicator variables for whether the individual is black, Asian, or Native American, female, married, has children of one's own,[13] and a categorical variable for one's age in 3-to-5-year increments. We also control for the seasonally-adjusted unemployment rate in her state during that month, as a control for local economic conditions.[14] Finally, we include indicator variables for the interview wave, to control for the increased probability of non-random attrition in later waves. Sample means are reported in Appendix Table 1.

We choose a linear model for simplicity, but it is important to note that our results hold for a probit model. Using a linear model allows for easier interpretation of coefficients and marginal effects, and with our multiple methods of decomposition, we should attempt to keep the interpretation as simple, and as comparable between methods, as possible. While a Blinder-Oaxaca decomposition from a nonlinear regression is possible (Fairlie 2005), the calculation of the trend decomposition becomes unwieldy when adding the normal distribution, and is beyond the scope of this paper.[15]

We use two different methods to determine how differences in particular variables (or categories of variables) between Hispanics and non-Hispanic Whites may contribute to the growing gap in the probability of being uninsured between the two groups: (1) the statistical significance of the trend line in the uninsured rate gap, regression-adjusted for each variable sequentially, and (2) pairwise Blinder-Oaxaca-style decomposition of the annual gap, and the change

[13] There was no difference when we included the number of children as a categorical variable (1, 2, 3 or more). Also, our coefficients remained unchanged when including the interaction between Hispanic and having one's own children.

[14] Cawley and Simon (2005) find that state unemployment rate is positively correlated with insurance coverage, and that this correlation survives even after controlling for employment transitions.

[15] Of the 7.8 million person-month observations in our OLS regression, the predicted value of the dependent variable is below zero for 740,000 of them (9.4 percent), but 75 percent of those are between zero and -0.04. Only 78 observations have a predicted value of greater than one.

in the gap between 1984 and each year thereafter.[16]

5.1 Method 1 - Regression-Adjusted Trend Line

In the first method, we see how $\beta_1 + \delta_y$, the adjusted gap, changes when we add variables sequentially to the regression. After adding each group of related variables, we then determine whether the slope of the trend in the annual uninsured rate gaps is significantly different from zero.

We begin with the unadjusted trend line, the difference in the (unweighted)[17] uninsured rate between Hispanics and non-Hispanic Whites, averaged over each month in the specific calendar year.

We then add categories of variables one-by-one. If the variables were independent, then the order they are added to our regression would be immaterial. Of course, this is not the case. We must then determine an order for their addition. We choose to add variables in rough order of when they were endowed to the individuals in our sample. We start with the variables that are largely out of the individual's control: gender, race, age, and family structure. Next, individuals are either born or choose to live in the United States and, if the latter, whether to become citizens. Then, they decide what level of education to acquire. Next, they (and their families) settle on a certain living standard, which either meets or fails to meet the poverty threshold calculated by the U.S. government. Finally, the individual and his or her spouse choose how much to work, in what industry, and for what pay.

Our first adjusted trend line, therefore, includes controls for Hispanic, the variables in X_{it}, and the year and Hispanic-year dummy variables.

Second, we add citizenship variables. Each SIPP respondent is asked their citizenship status, country of birth, and year entering the U.S. (if non-native) once during their time in the panel, in a special "topical module" separate from the usual ("core") questions. The topical module including the immigration and citizenship questions was during the second interview wave for all panels but the 1984 (wave 8) and 1985 (wave 4) panels. Because these questions are asked just once, and early in the life of the panel, there are many individuals who are never asked about their citizenship status. Also, not surprisingly, there are also many individuals who have no response for these questions. We may worry that the late-arriving and/or non-responding individuals may not be randomly distributed with respect to the other variables, so we include "no answer" as

[16] The results of a decomposition of each variable's contribution to the change in the gap's trend in the style of Bound and Freeman (1992) are similar to both of the included methods.

[17] As the uninsured are more likely to leave the sample, and perhaps more so if they are also Hispanic, our unweighted estimates may actually underestimate the gap. Our results should thus be considered a lower bound on Hispanic-White coverage differences.

one possible citizenship status, along with naturalized citizen, noncitizen, and native-born (the omitted condition). We also interact citizenship status with Hispanic origin. In addition, we include a categorical variable for how long the individual has been in the U.S.: less than ten years, ten to twenty years, or twenty-plus years, with native-born (or missing information) as the omitted condition.[18]

Next, we add a categorical variable for one's education level: less than high school, high school degree only, or some college, with bachelors degree or higher as the omitted condition.[19]

We next add the proxy variable for public coverage eligibility, an indicator variable for whether family income is below the federal poverty line for that year. We include interactions between this proxy variable and Hispanic ethnicity, and between the proxy and the female indicator. In addition, we include interactions with each of the three citizenship indicators, and with both citizenship and Hispanic ethnicity.[20]

Finally, we add variables that account for labor supply and income. We add a categorical variable for the respondent's family income as a percent of the federal poverty level. We include controls for whether the individual is working full time, part time, self employed, or unemployed (the omitted condition), and a broad categorical variable for her industry, based on the job for which she earns the highest hourly wage (imputed or directly reported).[21] In addition, we include indicator variables for spouse's work status, on the theory that a spouse working full time should increase one's access to health insurance.

As we add more variables, our "adjusted gap" shrinks toward zero in each year. We measure the amount of the gap that remains unexplained in each year, and its significance, as the difference between the sum $\beta_1 + \delta_y$ (for year y) and zero. More importantly, as a measure of how much the upward trend in the gap remains unexplained, we run a linear trend through the unexplained portion of

[18] We have experimented with adding each of these variables and interactions separately. After adding citizenship status, the interactions and time in U.S. variables have little effect on the Hispanic coefficients; most are significant, though, so we include them anyway.

[19] We also tried interacting educational attainment with citizenship status; the coefficients on the interaction were mostly insignificant and changed the other coefficients negligibly, so the interactions were excluded from further analysis.

[20] We have also experimented with including interactions between the eligibility proxy and the year dummies, on the theory that policy changes have an impact on eligibility differently in some years. With these variables, we can explain a bit more of the gap in individual years in the 1980's, but somewhat less in the later years, so the adjusted line actually gets steeper (i.e, we explain less of the upward trend). Therefore, we have not included the interactions in our final model.

[21] We can also include an indicator variable for union membership. If we do, though, we lose all observations in 1983 and most of 1984, because union status was not part of the SIPP in the 1984 panel. As the results including a union dummy are not substantially different, we have opted to restore the full sample and drop the union dummy from our regressions.

the gap in each year and determine its significance.[22]

5.2 Method 2 - Pairwise Blinder-Oaxaca Change Decompositions

As mentioned above, conclusions made about the analysis of Method 1 are somewhat limited, because the order that the variables are added to the regression controls the amount of influence we detect. For instance, if income has a large effect on the Hispanic-year coefficients, but much of that effect is sopped up by controlling for variables that are added earlier, such as education or public coverage eligibility, we may incorrectly conclude that income is not an important explanatory variable.

A more thorough method, and one independent of the order of addition for variables, is the Blinder-Oaxaca decomposition. In this paper, we perform two different types of Blinder-Oaxaca decomposition. First, we decompose the contribution of particular variables to the annual uninsured rate gap between Hispanics and non-Hispanic Whites. Then, we decompose the contribution of each variable to the change in the gap between the base year and each given year.

Adapting the setups from Blinder (1973) and Oaxaca (1973), we start with two mutually exclusive and exhaustive groups, non-Hispanic Whites (denoted W) and Hispanics (H). For a particular outcome y (in our primary estimates, the indicator variable for uninsured), we have two parallel models of the effect of variables X on that outcome y:

$$\begin{aligned} y_H &= \beta_H X_H + \epsilon_H, \\ y_W &= \beta_W X_W + \epsilon_W, \end{aligned} \tag{2}$$

where the error terms ϵ_H and ϵ_W are mean zero. We expect $\hat{y}_H > \hat{y}_W$, where $\hat{y}_J = \hat{\beta}_J \overline{X}_J$, the fitted value for y for group $J \in \{H, W\}$. The difference between the two fitted outcomes (suppressing the hat notation for simplicity) is:

$$y_H - y_W = \beta_H \overline{X}_H - \beta_W \overline{X}_W. \tag{3}$$

Traditional Blinder-Oaxaca decomposition then adds and subtracts either $\beta_W \overline{X}_H$ or $\beta_H \overline{X}_W$. In the former case, after grouping terms, the following equation results:

$$y_H - y_W = \beta_W (\overline{X}_H - \overline{X}_W) + (\beta_H - \beta_W)\overline{X}_H. \tag{4}$$

which decomposes the difference between y in populations H and W into the portion that can be explained by differences in the mean of the variables X in the two groups (the first part, commonly called the “explained” portion) and

[22]Results using a quadratic trend are similar.

the portion owing to differences in the coefficients between the two groups for the same values of X (the latter part, or the "unexplained" portion).[23]

A similar process can be used to decompose the *change* in the uninsured rate gap between Hispanics and white non-Hispanics between any two points in time (Altonji and Blank, 1999). For members of group $J \in \{H, W\}$, the change in y between periods t and t' can be written as

$$\begin{aligned}
\Delta y_J \equiv y_{Jt'} - y_{Jt} &= \beta_{Jt'}\overline{X}_{Jt'} - \beta_{Jt}\overline{X}_{Jt} \\
&= \beta_{Jt'}\overline{X}_{Jt'} - \beta_{Jt}\overline{X}_{Jt} + \beta_{Jt'}\overline{X}_{Jt} - \beta_{Jt'}\overline{X}_{Jt} \\
&= \beta_{Jt'}(\overline{X}_{Jt'} - \overline{X}_{Jt}) + (\beta_{Jt'} - \beta_{Jt})\overline{X}_{Jt},
\end{aligned} \tag{7}$$

where the first line uses the definition $\Delta z_J = z_{Jt'} - z_{Jt}$, for some vector z.

Then, we can decompose the change in the gap between H and W for outcome y as

$$\begin{aligned}
\Delta y_H - \Delta y_W = &\beta_{Ht'}(\overline{X}_{Ht'} - \overline{X}_{Ht}) + (\beta_{Ht'} - \beta_{Ht})\overline{X}_{Ht} \\
&- \beta_{Wt'}(\overline{X}_{Wt'} - \overline{X}_{Wt}) - (\beta_{Wt'} - \beta_{Wt})\overline{X}_{Wt}.
\end{aligned} \tag{8}$$

Again, we use the white non-Hispanic coefficient $\beta_{Wt'}$ and β_{Wt} as benchmarks.[24]

---

[23] Alternatively, if we added and subtracted $\beta_W\overline{X}_H$ from equation (5), our decomposition equation would be

$$y_H - y_W = \beta_H(\overline{X}_H - \overline{X}_W) + (\beta_H - \beta_W)\overline{X}_W. \tag{5}$$

Oaxaca and Ransom (1994) demonstrate that these two methods bound the actual Hispanic effect on the probability of being uninsured. Jann (2005) puts the Blinder-Oaxaca decomposition in general form:

$$y_H - y_W = \beta^\star(\overline{X}_H - \overline{X}_W) + [(\beta_H - \beta^\star)\overline{X}_H + (\beta^\star - \beta_W)\overline{X}_W], \tag{6}$$

where $\beta^\star$ is the "benchmark coefficient." In the case outlined in the text, $\beta^\star = \beta_W$; in the footnote case, $\beta^\star = \beta_H$. Reimers (1983) suggests using the arithmetic average of the two coefficients $\beta^\star = \frac{1}{2}\beta_H + \frac{1}{2}\beta_W$. Cotton (1988) prefers the weighted average, $\beta^\star = \ell_H\beta_H + (1 - \ell_H)\beta_W$, where ℓ_H is the percent of the total population in group H. Finally, Neumark (1988) uses the coefficients from a regression run on the pooled sample (including both Hispanics and white non-Hispanics, in our case) as $\beta^\star$. Oaxaca and Ransom find that the Cotton or Neumark decompositions have the smallest standard errors, and are thus preferred to more extreme benchmarks like the Hispanic coefficient. We will present results from only the decomposition that uses the white non-Hispanic coefficient as the benchmark ($\beta^\star = \beta_W$), since most of our sample is white non-Hispanic (between 84 and 94 percent) and therefore the population-weighted average coefficients and the pooled coefficients are very similar to β_W.

[24] The general form for the decomposition in the style of Jann (2005) is

$$\begin{aligned}
\Delta y_H - \Delta y_W = &\beta^\star_t[(\overline{X}_{Ht'} - \overline{X}_{Wt'}) - (\overline{X}_{Ht} - \overline{X}_{Wt})] \\
&+ (\beta^\star_{t'} - \beta^\star_t)(\overline{X}_{Ht'} - \overline{X}_{Wt'}) \\
&+ [(\beta_{Ht'} - \beta^\star_{t'}) - (\beta_{Ht} - \beta^\star_t)]\overline{X}_{Ht} + [(\beta^\star_{t'} - \beta_{Wt'}) - (\beta^\star_t - \beta_{Wt})]\overline{X}_{Wt} \\
&+ (\beta_{Ht'} - \beta^\star_{t'})(\overline{X}_{Ht'} - \overline{X}_{Ht}) + (\beta^\star_{t'} - \beta_{Wt'})(\overline{X}_{Wt'} - \overline{X}_{Wt}).
\end{aligned}$$

So the decomposition between our base year and each given year is

$$\begin{aligned}\Delta y_H - \Delta y_W = {} & \beta_{Wt}[(\overline{X}_{Ht'} - \overline{X}_{Wt'}) - (\overline{X}_{Ht} - \overline{X}_{Wt})] \\ & + (\beta_{Wt'} - \beta_{Wt})(\overline{X}_{Ht'} - \overline{X}_{Wt'}) \\ & + [(\beta_{Ht'} - \beta_{Wt'}) - (\beta_{Ht} - \beta_{Wt})]\overline{X}_{Ht} \\ & + (\beta_{Ht'} - \beta_{Wt'})(\overline{X}_{Ht'} - \overline{X}_{Ht}). \end{aligned} \tag{9}$$

Altonji and Blank (1999) identify the first line as the effect of relative changes over time in the observed characteristics of the two groups, and the second line as the effect of changes over time in the white non-Hispanic coefficient, holding differences in the observed characteristics fixed; these two parts make up the "explained" portion of the change decomposition. The third line is the effect of changes over time in the relative coefficients between the two groups, and the fourth line "captures the fact that changes over time in the characteristics of [the two groups] alter the consequences of differences in group coefficients" (p. 3226); these two lines make up the "unexplained" portion of the change decomposition.[25] We will display only the total explained and unexplained portions of equation (9), rather than separately present the results for each of the four portions.

We will apply the change decomposition to each year from 1985 to 2003 as it relates to 1984,[26] the year when the average uninsured rates for Hispanics (27.5 percent) and white non-Hispanics (15.3 percent) were closest. In terms of equation (9), t will be 1984 in each case, while we take each subsequent year, where the change in the gap from 1984 is positive, and progressively getting larger as we advance through our sample, as t'.

One disadvantage of this method is that we are unable to ascertain the amount of the annual gap, or the change in the annual gap between a given year and 1984, that can be explained by variables that are perfectly collinear with Hispanic. In particular, we can no longer include the interaction between Hispanic and the public coverage eligibility proxy, or Hispanic and citizenship status. While the coefficients on these variables (in the last regression of Method 1) are significantly different from zero, they are of relatively small magnitude, and their exclusion leaves the key coefficients largely unchanged.

Provided that these excluded variables are relatively unimportant, we should expect, for each year y, $\beta_1 + \delta_y$ from Method 1 to be approximately equal to the unexplained portion of the annual gap for that year from Method 2. If, in addition, the linear trend we draw through the all-inclusive regression adjusted trend line from Method 1 is a good approximation, then the slope of this line

[25] Fairlie and Sundstrom (1999) have a mathematically equivalent change decomposition, but have a broader definition of the "explained" portion that would include parts of the third line of equation (9).

[26] We skip the change from 1983 to 1984 in this analysis because the gap decreased between these two years.

should match the unexplained portion of the change between 1984 and 2003 from Method 2.

6 Results

6.1 Regression Adjusted Trend Line

The first column of Table 2 is the actual (unweighted, because the regressions are all unweighted) gap in uninsured rates between Hispanics and non-Hispanic Whites, averaged over the months in each year. The actual gap grew from a low of 12.4 in 1984 to a high of 28.8 in 2001, decreasing slightly in the last two years of the sample.[27] Fitting a linear trend to the numbers in this column (second panel), we find a slope of 0.644, and we can reject the null of a zero slope at 99 percent confidence level.

The remaining columns of Table 2 present the adjusted Hispanic-White gap, the sum of the coefficient on the Hispanic intercept term plus each Hispanic-year coefficient ($\beta_1 + \delta_y$ from equation 1). In the second column, we add the X variables: gender, race and ethnicity, family structure, age categories, and the state unemployment rate. On average, the annual gaps shrink by three percent, or about 0.7 percentage points. Each gap is at least a little smaller, with the 1983 gap shrinking by the greatest percent (5.8 percent, or 0.9 percentage points). The slope of the fitted linear trend line is flatter at 0.606 percentage points per year, a decrease of 5.8 percent in the slope, but is still significantly different from zero.

Next, we add citizenship variables, including indicators for naturalized citizen and non-citizen, an indicator for no citizenship information, interactions of each of those three statuses with the Hispanic dummy, and a categorical variable for how long the individual has spent in the United States. On average, the annual gaps fall by an additional nine percentage points, or about 46 percent. The reduction is largest for the early years; the gap shrunk by at least 50 percent in eight out of the first thirteen years, but by less than 40 percent in the 2000's. The trend line is 22.5 percent flatter than the unadjusted trend, but is still significantly different from zero.

In the fourth column, we add categorical variables for educational attainment. The annual gaps are an additional 3.7 percentage points smaller on average, and the average adjusted gap is now a third of the unadjusted gap for that year. Still, this reduction is substantially smaller for the more recent years in the sample, and the linear trend still has a statistically significant slope of 0.4 percentage points per year.

[27]The slight decrease in the gap exists even when we reweight to correct for sampling attrition.

Table 2: Predicted Hispanic - White Non-Hispanic Gaps in Uninsured Rate (Method 1)

	Actual Gap (Percentage Points)	Predicted Gap (Percentage Points)				
		Add Basic Controls	Add Citizenship and Time in U.S.	Add Education	Add Public Coverage Eligibility Proxy and Interactions	Add Income, Work Status, and Industry
1983	15.8	14.9	7.4	4.1	3.8	2
1984	12.4	12.1	4.5	1.3	1.7	-0.1
1985	14.0	13.6	6.1	3	3.9	1.7
1986	18.9	18.3	11	7.9	8.3	5.7
1987	18.4	17.9	10.5	7.3	7.6	5.3
1988	18.5	18.1	10.1	7.1	7.4	5.1
1989	18.9	18.7	10	6.8	7.2	5.1
1990	20.4	20.3	10.4	7	7.1	5.4
1991	18.8	18.5	8.6	5.2	5.1	3.4
1992	18.5	18	8.1	4.8	4.7	3
1993	18.5	17.9	8	4.6	4.4	2.7
1994	18.3	17.6	8	4.7	4.7	3
1995	18.7	18	8.9	5.4	5.5	3.5
1996	24.9	23.7	13.6	9.5	9.8	7.6
1997	23.7	22.7	13	8.7	9.2	7
1998	23.6	22.6	13	8.7	9.2	7.1
1999	23.1	22.2	12.8	8.6	9.1	7.5
2000	28.0	26.7	17.4	12.6	13	10.7
2001	28.8	27.6	17.9	13.1	14.2	12.3
2002	27.0	26	16.8	12.2	13.3	11.6
2003	25.5	24.7	15.8	11.3	12.5	10.8
	Linear Trend, 1983-2003					
Coefficient	0.644	0.606	0.499	0.409	0.444	0.455
t-statistic	8.7	8.6	6.4	5.6	5.6	6.2
Percent of Trend in Actual Gap Explained		5.8%	22.5%	36.5%	31.1%	29.3%
	Linear Trend, 1983-1995					
Coefficient	0.364	0.362	0.120	0.097	0.061	0.093
t-statistic	2.7	2.6	0.9	0.7	0.4	0.7
Percent of Trend in Actual Gap Explained		0.6%	66.9%	73.4%	83.2%	74.5%
	Linear Trend, 1996-2003					
Coefficient	0.491	0.512	0.639	0.563	0.694	0.764
t-statistic	1.6	1.9	2.6	2.5	3.1	3.8
Percent of Trend in Actual Gap Explained		-4.2%	-30.1%	-14.6%	-41.3%	-55.6%

We next add our proxy variable for public coverage eligibility, plus interactions between the proxy and Hispanic, female, each citizenship status, and both Hispanic and citizenship status. Accounting for public coverage eligibility actually increases the gap slightly in most years, particularly in the years after welfare reform. This leads to a steeper slope in the trend line, though still a 31 percent slower average annual increase in the gap than the unadjusted figures.

Finally, we add income and labor supply variables, including indicators for whether one is full time, part time, or self employed, the same employment status categories for one's spouse, and five industry categories. Controlling for these variables reduces each of the gaps, undoing the reversal caused by adding public coverage eligibility. Again, though, the reduction in the gap is largest in the early years; in fact, the 1984 gap disappears. Still, almost half of the gap in each year of the 2000's remains unexplained. As a result, the trend line gets even steeper, with the gap growing at a statistically significant 0.455 percentage points per year, only 29.3 percent less steep than the unadjusted trend. Despite significant contributions by citizenship and education, about 70 percent of the divergence in the uninsured rate between Hispanics and non-Hispanic Whites remains unexplained.

Figure 9 graphs these adjusted gaps. It is clear from the graph that citizenship and education have the largest effects, while income also reduces the annual gaps, but most of its effect worked through variables added previously. Drawing one trend line through all of the points in each series, we see that we can explain some of the growth of the gap, but that a significant upward trend remains even in the all-inclusive series.
In Figure 9, there appears to be a break at 1996, possibly due to subtle differences between earlier panels and the 1996 panel (though most of the variables we used appear to be largely unchanged). It may be more appropriate to draw two trends lines, one from 1983 to 1995 and another from 1996 to the end of the sample. The early trend line (1983-1995, third panel of Table 2) has a positive slope of 0.36 percentage points per year, significantly different from zero, for the unadjusted gap and the gap adjusted for only the X variables, but becomes essentially flat once we control for citizenship. The later trend line (1996-2003, bottom panel of Table 2), though, is actually significantly different from zero only at the 90 percent confidence interval for the unadjusted gap, but with each additional group of control variables gets progressively steeper; the all-inclusive line has a slope of 0.76 percentage points per year, and we reject the null hypothesis of a zero slope.

The other coefficients are largely in the expected direction, and because of our large sample size (7.8 million person-months, with standard errors clustered by 312,000 individuals), almost all statistically significant. Females and married individuals are less likely to be uninsured, while those with children are more likely to be uninsured. The uninsured rate is decreasing in age (up through age 55), education, and family income, and increasing in the state unemployment

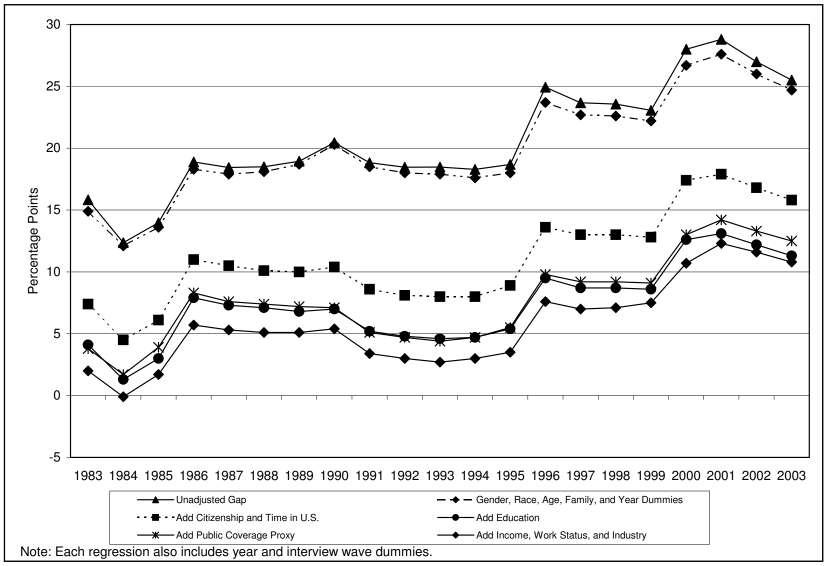


Figure 9: Hispanic - White Non-Hispanic Gap in Uninsured Rate

rate. Non-citizens are most likely to be uninsured, more so if they are also Hispanic, and while naturalized citizens are more likely to be uninsured than native-born citizens, there is no difference between Hispanic and non-Hispanic naturalized citizens. Foreign-born individuals that have been in the U.S. for more than ten years are less likely to be uninsured than even the native-born. Those with family income less than the poverty level (our proxy for public coverage eligibility) are far more likely to be uninsured, though being female, Hispanic, or a non-Hispanic immigrant reduces this probability. Hispanic immigrants (both non-citizens and naturalized citizens) are particularly vulnerable to being uninsured. Finally, full time workers, or those with spouse's working full time, are least likely to be uninsured, while the self-employed are actually worse off than the unemployed, though a self-employed spouse is better than an unemployed spouse. Coefficients are unchanged, for the most part, when adding each new set of controls, except for those variables that are directly correlated, such as citizenship variables interacted with the public coverage eligibility proxy, or the married dummy with spouse's work status.

6.2 Blinder-Oaxaca Decompositions

In table 3, we present the results from the Blinder-Oaxaca decompositions of the Hispanic-White uninsured rate gap in each year. The actual gap is listed in the left-most column, the middle section lists the total percentage point "explained" effect from each variable and the "unexplained" portion, and the section to the right calculates how much each variable "explains" of the actual gap.

Except for citizenship in 1984, each variable contributes some positive amount to "explaining" the actual gap in all years. Differences in educational attainment consistently explain between ten and 20 percent of the actual gap, growing slightly more important in later years. Income and labor market differences seem to be important from 1990 through 1995, but otherwise do not contribute much. Citizenship becomes much more important starting in 1990, but remains less influential than differences in education level. Differential public coverage eligibility is crucial from 1984 to 1989, explaining up to half of the actual gap, but gradually fades in importance. Still, in most years, the largest portion of the actual gap is "unexplained," growing to more than half of the actual gap in the 2000's.

Tables 4 and 5 display the results of the Blinder-Oaxaca decomposition of the change in the uninsured rate gap between a base year and each given year. In Table 4, the base year is 1984, when the gap was the smallest. We focus on the last line, the decomposition of the change in the gap between 1984 and 2003. We find that, much like with Method 1, education and citizenship contribute the most to the "explained" portion of the actual growth of the uninsured rate gap. Of the 13.1 percentage point increase from 1984 to 2003, differences between Hispanics and non-Hispanic Whites in education account for 3.2 percentage points (24.1 percent). Citizenship accounts for an almost identical amount, with income somewhat less important. Much as we saw (though to a lesser extent) in Method 1, Hispanic-White differences in public coverage eligibility actually work in the wrong direction, suggesting that Hispanics' greater likelihood of living below the poverty line should actually make the increase in the gap greater than it is.

Again, the largest portion of the decomposition in Table 4 is the "unexplained" part. Nearly ten percentage points, or 73.3 percent, of the 13.1 percentage point increase in the uninsured rate gap from 1984 to 2003 remains unexplained. The relative importance of the unexplained portion is consistent across endpoints, but generally larger for the changes between 1984 and any of the years in the 2000's than for the changes through earlier years.

In Figure 9, we saw that two trends lines, before and after 1996, were a better fit to the adjusted gaps. We can easily repeat the same exercise with Method 2. The change between 1984 and 1995 is displayed the shaded line near the middle of Table 4. Citizenship and income differences between Hispanics and Whites seem to be about equally important, with education contributes slightly less. Here, unlike Method 1, a substantial portion of the growth in the gap, more than 60 percent, remains unexplained.

Table 5 repeats the exercise with 1996 as the base year, similar to the later trend line in Method 1. The uninsured rate gap in 2003 is only 0.6 percentage points higher than 1996, but Hispanic-White citizenship and public coverage eligibility differences suggest that the gap should be larger. As a result, the

Table 3: Blinder-Oaxaca Decompositions for Annual Hispanic-White Non-Hispanic Uninsured Rate Gap

	Actual	Percentage Points						Percent of Actual Gap					
	Gap	Inc	Educ	Cit	Pub	Oth	Unexp	Inc	Educ	Cit	Pub	Oth	Unexp
1983	15.8	1.1	2.1	0.1	5.7	1.5	5.4	7.2	13.1	0.4	36.0	9.3	33.9
1984	12.4	0.2	1.7	-0.2	7.1	0.5	3.0	1.9	13.6	-1.3	57.8	4.2	23.9
1985	14.0	0.8	1.8	0.5	7.0	0.5	3.4	5.4	13.1	3.5	49.8	3.7	24.4
1986	18.9	0.8	2.0	1.2	7.2	0.7	7.0	4.4	10.4	6.3	38.1	3.9	36.9
1987	18.4	0.4	1.8	0.9	7.7	0.7	6.9	2.4	9.5	4.9	41.6	4.0	37.6
1988	18.5	0.7	1.7	0.2	7.8	0.7	7.4	3.7	9.0	1.3	42.0	4.0	40.0
1989	18.9	0.8	2.0	1.1	7.2	1.0	6.8	4.4	10.6	5.9	38.0	5.1	36.0
1990	20.4	2.8	3.1	2.9	3.6	0.6	7.4	13.9	15.1	14.2	17.5	3.0	36.4
1991	18.8	2.5	2.8	2.0	4.2	0.4	6.9	13.0	15.0	10.5	22.6	2.2	36.7
1992	18.5	2.3	3.0	1.5	3.6	0.5	7.5	12.5	16.4	8.0	19.8	2.8	40.6
1993	18.5	2.5	3.2	1.5	3.9	0.8	6.7	13.3	17.1	8.0	21.0	4.4	36.1
1994	18.3	2.5	2.8	1.1	2.8	1.4	7.7	13.7	15.5	6.0	15.1	7.4	42.2
1995	18.7	2.2	3.0	1.7	3.6	1.4	6.8	11.6	16.1	9.3	19.1	7.5	36.4
1996	24.9	1.8	4.0	3.4	4.1	1.5	10.2	7.1	16.1	13.8	16.4	5.9	40.7
1997	23.7	1.9	3.5	2.3	4.2	1.4	10.3	8.2	15.0	9.8	17.6	5.8	43.6
1998	23.6	1.8	3.7	2.3	3.7	1.6	10.5	7.4	15.7	9.8	15.7	6.8	44.6
1999	23.1	1.5	3.7	2.5	3.2	1.4	10.6	6.6	16.2	10.8	14.0	6.2	46.2
2000	28.0	2.3	4.2	1.5	3.4	1.6	15.1	8.2	14.9	5.3	12.0	5.7	53.8
2001	28.8	2.3	4.5	3.1	2.2	1.4	15.3	7.9	15.6	10.8	7.6	4.9	53.2
2002	27.0	2.2	5.1	2.9	1.9	1.3	13.5	8.3	18.9	10.7	7.1	4.9	50.0
2003	25.5	2.3	4.8	3.0	1.8	0.9	12.6	9.1	19.0	11.7	7.2	3.7	49.4
Average	20.7	1.7	3.1	1.7	4.6	1.0	8.6	8.2	14.9	8.2	22.0	5.1	41.6

Note: Benchmark coefficient is from the White Non-Hispanic regression.
Key:
Inc includes family income as a percent of poverty, work status (own and spouse), and industry of employment.
Educ includes the three categories of educational attainment.
Cit includes citizenship status and time lived in the U.S.
Pub includes the public coverage eligibility proxy and its interaction with gender and citizenship status.
Oth includes gender, marital status, presence of children, wave of interview, age, and state unemployment rate.
Unexp is the total of the "unexplained" portions of the decomposition, including the constant term.

Table 4: Blinder-Oaxaca Decompositions for the Change in the Hispanic-White Non-Hispanic Uninsured Rate Gap

Change between 1984 and...	Actual Change in the Gap	Percentage Point Change						Percent of Actual Change in Gap					
		Inc	Educ	Cit	Pub	Oth	Unexp	Inc	Educ	Cit	Pub	Oth	Unexp
1985	1.6	0.5	0.2	0.7	-0.2	0.0	0.5	32.8	9.5	40.5	-11.1	0.1	28.3
1986	6.5	0.6	0.3	1.4	0.0	0.2	4.0	9.1	4.4	20.7	0.7	3.4	61.7
1987	6.1	0.2	0.1	1.1	0.5	0.2	4.0	3.4	1.3	17.4	8.8	3.7	65.4
1988	6.1	0.5	0.0	0.4	0.6	0.2	4.4	7.6	-0.1	6.7	10.0	3.6	72.3
1989	6.6	0.6	0.3	1.3	0.1	0.5	3.9	9.1	5.1	19.4	0.8	6.9	58.7
1990	8.1	2.6	1.4	3.1	-3.6	0.1	4.5	32.2	17.4	37.8	-44.1	1.2	55.4
1991	6.5	2.2	1.1	2.1	-2.9	-0.1	3.9	34.4	17.7	33.2	-44.8	-1.5	61.1
1992	6.1	2.1	1.3	1.6	-3.5	0.0	4.5	34.0	22.0	26.9	-57.4	-0.1	74.5
1993	6.1	2.2	1.5	1.6	-3.3	0.3	3.7	36.6	24.3	26.7	-53.4	5.0	60.8
1994	5.9	2.3	1.2	1.3	-4.4	0.8	4.8	38.6	19.5	21.4	-74.1	14.3	80.3
1995	6.3	1.9	1.3	1.9	-3.6	0.9	3.8	30.6	21.1	30.1	-56.6	13.9	60.8
1996	12.6	1.5	2.3	3.6	-3.1	1.0	7.2	12.2	18.5	28.6	-24.3	7.6	57.3
1997	11.3	1.7	1.9	2.5	-3.0	0.9	7.4	15.1	16.5	21.9	-26.4	7.7	65.2
1998	11.2	1.5	2.0	2.5	-3.5	1.1	7.6	13.6	18.0	22.1	-30.9	9.7	67.5
1999	10.7	1.3	2.1	2.6	-3.9	0.9	7.7	12.2	19.2	24.7	-36.5	8.5	71.9
2000	15.6	2.1	2.5	1.7	-3.8	1.1	12.1	13.2	16.0	10.6	-24.2	7.0	77.5
2001	16.4	2.0	2.8	3.3	-5.0	0.9	12.4	12.4	17.2	20.0	-30.2	5.4	75.3
2002	14.6	2.0	3.4	3.1	-5.2	0.8	10.5	13.7	23.4	20.9	-35.7	5.6	72.1
2003	13.1	2.1	3.2	3.1	-5.3	0.4	9.6	15.9	24.1	23.9	-40.4	3.2	73.3
Average	9.0	1.6	1.5	2.0	-2.8	0.5	6.1	17.5	16.8	22.6	-30.8	6.0	67.9

Note: Benchmark coefficient is from the White Non-Hispanic regression.
Key:
Inc includes family income as a percent of poverty, work status (own and spouse), and industry of employment.
Educ includes the three categories of educational attainment.
Cit includes citizenship status and time lived in the U.S.
Pub includes the public coverage eligibility proxy and its interaction with gender and citizenship status.
Oth includes gender, marital status, presence of children, wave of interview, age, and state unemployment rate.
Unexp is the total of the "unexplained" portions of the decomposition, including the constant term.

"unexplained" portion is larger than the true change in the gap between 1996 and each subsequent endpoint year, a result that's analogous to the adjusted gap trend line for the later years in Method 1 getting steeper with each additional control.

Table 5: Blinder-Oaxaca Decompositions for the Change in the Hispanic-White Non-Hispanic Uninsured Rate Gap

Change between 1996 and...	Actual Change in the Gap	Percentage Point Change						Percent of Actual Change in Gap					
		Inc	Educ	Cit	Pub	Oth	Unexp	Inc	Educ	Cit	Pub	Oth	Unexp
1997	-1.3	0.2	-0.5	-1.1	0.1	-0.1	0.2	-13.7	36.4	88.4	-5.3	7.1	-13.0
1998	-1.4	0.0	-0.3	-1.1	-0.4	0.1	0.4	0.9	22.6	82.0	29.2	-9.1	-25.6
1999	-1.9	-0.2	-0.3	-1.0	-0.9	0.0	0.5	12.7	14.5	51.0	45.6	2.3	-26.1
2000	3.1	0.5	0.2	-2.0	-0.7	0.1	4.9	17.0	5.6	-63.8	-23.6	4.5	160.3
2001	3.9	0.5	0.5	-0.3	-1.9	-0.1	5.2	12.8	12.8	-8.1	-49.5	-1.7	133.8
2002	2.1	0.5	1.1	-0.5	-2.2	-0.1	3.3	22.9	53.2	-26.4	-105.2	-7.0	162.4
2003	0.6	0.5	0.8	-0.5	-2.3	-0.5	2.4	95.3	146.4	-80.1	-393.2	-92.6	424.2
Average	0.7	0.3	0.2	-0.9	-1.2	-0.1	2.4	38.8	30.8	-128.1	-163.3	-12.1	334.0

Note: Benchmark coefficient is from the White Non-Hispanic regression.
Key:
Inc includes family income as a percent of poverty, work status (own and spouse), and industry of employment.
Educ includes the three categories of educational attainment.
Cit includes citizenship status and time lived in the U.S.
Pub includes the public coverage eligibility proxy and its interaction with gender and citizenship status.
Oth includes gender, marital status, presence of children, wave of interview, age, and state unemployment rate.
Unexp is the total of the "unexplained" portions of the decomposition, including the constant term.

6.3 Decompositions by Insurance Source

Figure 10 repeats the analysis from Method 1, with an indicator for public coverage, rather than uninsured, as the dependent variable. Adding our X variables reduces the average annual public coverage gap by 15 percent. Accounting for citizenship, the adjusted gap is actually larger than the unadjusted gap in all years. The education variables induce about a 30 percent reduction in the average gap and, not surprisingly, accounting for public coverage eligibility makes the gap nearly disappear. Adding variables does not change the shape of the public coverage gap's time series; the unadjusted gap and all of the adjusted gaps each initially decline, increase during the early 1990's, fall through 2000, and then rise again at the end of our sample. The public coverage gap between Hispanics and Whites is almost fully explained by Hispanics' greater probability of having family income below the poverty line, and unlike the uninsured rate gap, there does not seem to be a concern for the widening of the public coverage gap.

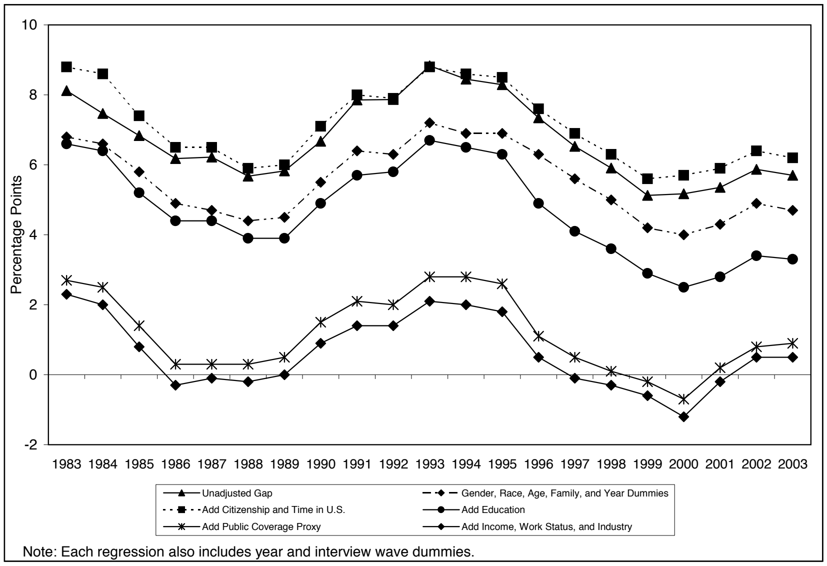


Figure 10: Hispanic - White Non-Hispanic Gap in Public Coverage Rate

If the public coverage gap can be almost completely explained away, but the uninsured rate gap persists even after controlling for observable characteristics, then the private coverage adjusted gap must also be growing, as we see in Figure 11. The unadjusted gap is negative, meaning the private coverage rate for Hispanics is lower than the rate for non-Hispanic Whites, and getting progressively more negative as our sample advances. Adding citizenship variables reduces the private coverage gap by about a third, and education, public coverage eligibility (which could have a crowdout effect on private coverage), and income each reduce the gap by about 15 percent. Still, about a quarter of the average annual gap remains unexplained, and the adjusted trend is significantly upward, whether we measure the trend over the entire sample or in two pieces before and after 1996.

The persistence of the private coverage gap between Hispanics and non-Hispanic Whites, even after accounting for observables, is not due to employer-sponsored coverage. Figure 12 shows that citizenship explains about half of the average annual gap in coverage from one's own employer, while the difference in educational attainment covers nearly the rest. While the own employer coverage gap is increasing at a small but significant rate, 0.16 percentage points per year, nearly all of this growth is in the last years of our sample; accounting for observables, particularly public coverage eligibility and income, actually makes the post-1996 growth rate in this gap larger.

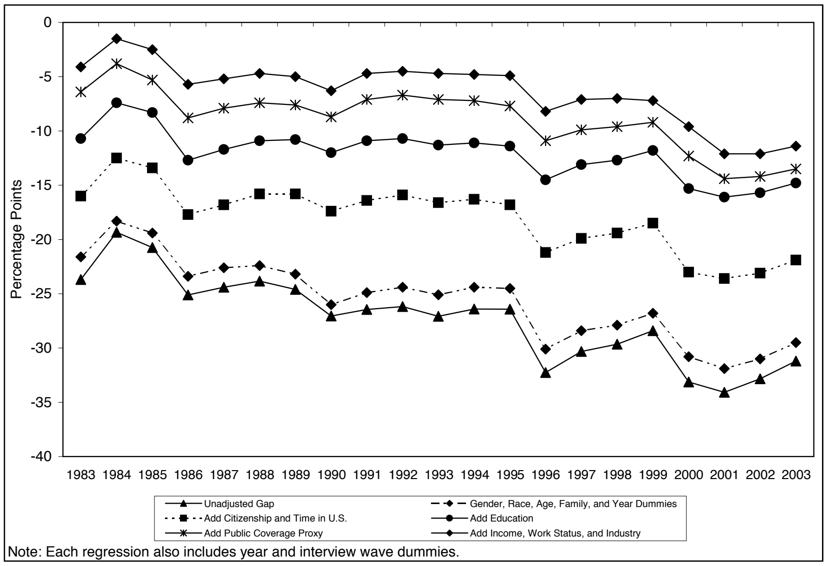


Figure 11: Hispanic - White Non-Hispanic Gap in Private Coverage Rate

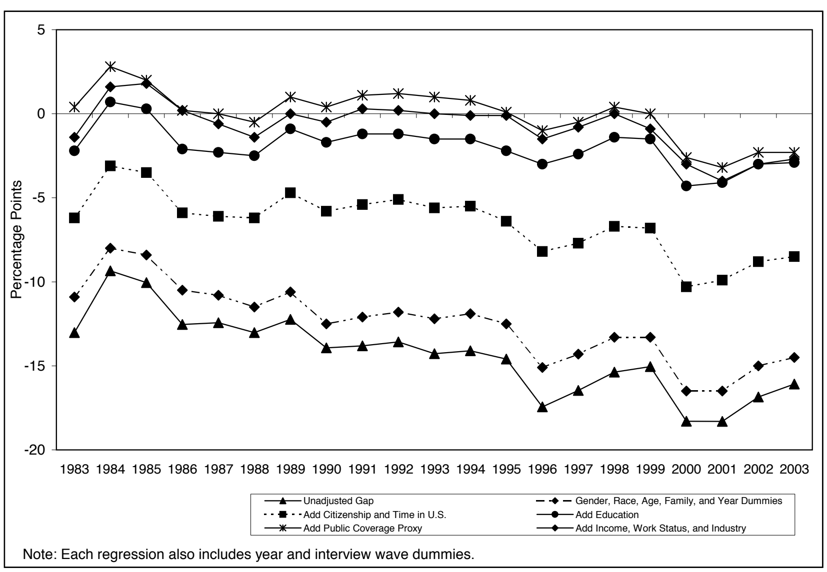


Figure 12: Hispanic - White Non-Hispanic Gap in Rate of Own Employer-Sponsored Coverage

The growing private coverage gap also does not seem to be due to a divergence in non-group coverage, as seen in Figure 13. The gap in own private coverage through a source other than one's employer is never larger than four percentage points, and accounting for observables reduces the gap to less than two percentage points. There is no significant widening of the gap; in fact, unlike most of the other graphs, the trend from 1996 on is toward zero, reflecting a tightening of the gap.

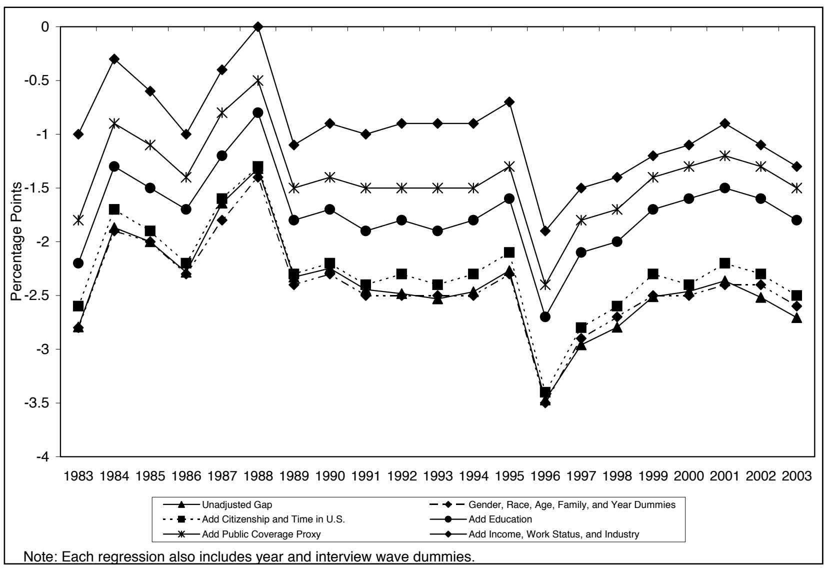


Figure 13: Hispanic - White Non-Hispanic Gap in Rate of Own Private non-ESI Coverage

The source of the widening adjusted gap in the uninsured rate seems to be due to the decreased likelihood of private coverage through someone else, usually a spouse or parent. While this rate has remained constant at around 28 percent for non-Hispanic Whites, the proportion of Hispanics covered by someone else fell from 21 percent to 14 percent (Figure 14). Citizenship explains only about 14 percent of the average annual gap, education another 8 percent, public coverage eligibility another 12 percent, and income about 24 percent, but more than 40 percent of the average annual gap remains unexplained. The adjusted trend line, including all variables, is just as steep as the unadjusted trend line, a significant increase of 0.2 percentage points per year, with the steepest portion after 1996.

6.4 Decompositions by Hispanic Subgroup

Figure 6, and previous studies, suggest that the uninsured rate has been much different for Mexicans living in the U.S. compared to other Hispanic subgroups. In Table 6, we present the results of separate decompositions for the growth

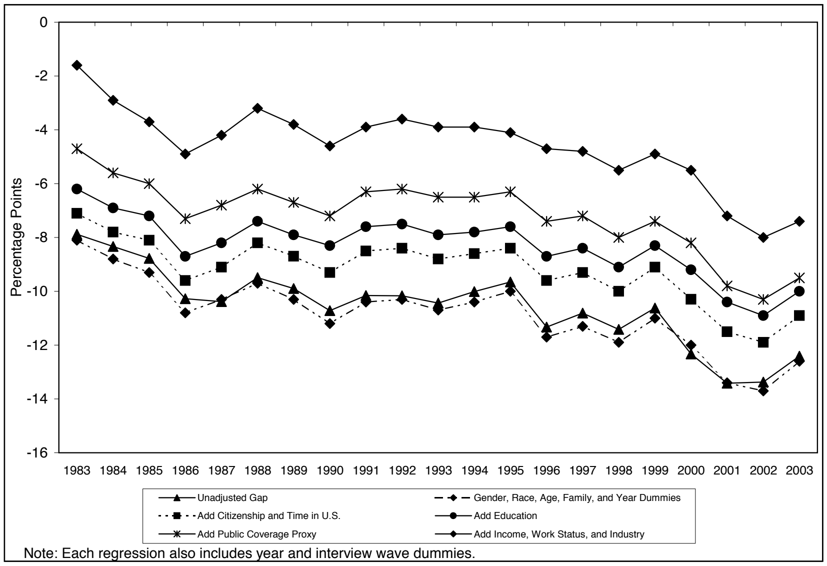


Figure 14: Hispanic - White Non-Hispanic Gap in Rate of Private Coverage From Someone Else

in the uninsured rate gap between each Hispanic subgroup and non-Hispanic Whites.

The largest subgroup, and the one with consistently the highest uninsured rate, is Mexicans, including both Mexican immigrants and U.S.-born Mexican-Americans. Between 1984 and 2003, the uninsured rate among Mexicans increased from 33.7 percent to 43.5 percent, an increase in the gap with non-Hispanic Whites of 14.7 percentage points. As we found in Table 4, differences in education and citizenship account for the largest portion of this gap, but 9.4 percentage points, or nearly two-thirds, of the Mexican-White gap remains unexplained. Similarly, the gap between Puerto Ricans and Whites, and other Hispanics (those who are neither Mexican, Puerto Rican, nor Cuban) and Whites, are largely left unexplained even after controlling for observables. The Cuban-White gap has actually shrunk slightly, despite differences in citizenship that suggest the gap should have widened.

We also repeat the analysis from Fronstin, Goldberg, and Robins (1997), decomposing the differences within the Hispanic group (Table 7). The uninsured rate among Mexicans increased much faster than the uninsured rate for Puerto Ricans or Cubans, but the widening of these intra-Hispanic gaps remains largely unexplained after controlling for observables. Again, education and citizenship differences between Mexicans and these two groups seem to be most important, especially considering nearly all individuals of Puerto Rican heritage are citizens. The gap between Mexicans and the remaining Hispanics did not widen nearly as much, and observables account for more than the actual gap.

Table 6: Blinder-Oaxaca Decompositions for the Change in Gap with Non-Hispanic Whites

	Actual Change in the Gap, 1984-2003	Percentage Point Change						Percent of Actual Change in Gap					
		Inc	Educ	Cit	Pub	Oth	Unexp	Inc	Educ	Cit	Pub	Oth	Unexp
Mexican	14.7	2.4	3.7	3.2	-4.5	0.4	9.4	16.5	25.4	21.7	-30.3	3.0	63.6
Puerto Rican	3.9	2.0	1.4	1.1	-11.4	0.6	10.2	51.8	36.8	27.1	-294.5	15.2	263.6
Cuban	-0.9	0.4	0.6	4.3	-4.7	0.0	-1.5	-42.4	-70.9	-476.1	522.0	2.9	164.4
Other Hispanic	11.6	1.2	2.3	3.6	-4.2	0.4	8.4	10.5	19.5	30.6	-36.3	3.2	72.5

Note: Benchmark coefficient is from the White Non-Hispanic regression.
"Other Hispanic" refers to Hispanics that are neither Mexican, Puerto Rican, nor Cuban.

Table 7: Blinder-Oaxaca Decompositions for the Change in Gap with Mexicans

	Actual Change in the Gap, 1984-2003	Percentage Point Change						Percent of Actual Change in Gap					
		Inc	Educ	Cit	Pub	Oth	Unexp	Inc	Educ	Cit	Pub	Oth	Unexp
Puerto Rican	-10.9	-1.6	-0.2	-3.7	1.0	1.1	-7.4	-14.8	-1.4	-34.1	8.8	10.1	-68.5
Cuban	-15.6	-3.9	3.8	4.1	-0.9	-3.9	-14.9	-24.8	24.5	26.5	-6.0	-24.9	-95.2
Other Hispanic	-3.1	-2.4	-1.7	-0.2	-0.5	0.8	1.0	-78.4	-55.4	-7.7	-15.9	24.3	33.1

Note: Benchmark coefficient is from the Mexican regression.
"Other Hispanic" refers to Hispanics that are neither Mexican, Puerto Rican, nor Cuban.
Key:
Inc includes family income as a percent of poverty, work status (own and spouse), and industry of employment.
Educ includes the three categories of educational attainment.
Cit includes citizenship status and time lived in the U.S.
Pub includes the public coverage eligibility proxy and its interaction with gender and citizenship status.
Oth includes gender, marital status, presence of children, wave of interview, age, and state unemployment rate.
Unexp is the total of the "unexplained" portions of the decomposition, including the constant term.

7 Discussion

Irrespective of the decomposition methodology, a large portion of the growth in the uninsured rate for Hispanics relative to white non-Hispanics remains unexplained, even after accounting for differences in income, education, citizenship, and other factors available in our data. To get an idea for the magnitude of this unexplained portion, we can do some simple accounting.

The uninsured rate for white non-Hispanic nonelderly adults fell from 19.5 percent in 1984 to 14.5 percent in 2003. The uninsured rate for Hispanics increased from 31.9 percent in 1984 to 40.0 percent in 2003, so the gap increased from 12.4 to 25.5 percentage points during that time. Of that 13.1 percentage point increase in the gap, we can explain 3.6 percentage points under either method.[28]

If the gap had remained constant at 1984 proportions, the Hispanic uninsured rate would have been 26.9 percent, or 6.4 million uninsured out of 24 million Hispanic nonelderly adults. Instead, there were 9.9 million Hispanic nonelderly adults without health insurance in the average month in 2003, an increase of 3.4 million over the counterfactual. Accounting for income, education, citizenship, and other observables, we can explain 920,000 (Method 2) to 930,000 (Method 1) of the difference between the actual and the counterfactual. In other words, 2.5 million of the increase in the number of uninsured Hispanic nonelderly adults remains unexplained by observables.

Another counterfactual can help us estimate how the growth of the Hispanic population in the U.S. has led to an increase in the number of uninsured. The non-Hispanic White population grew by 5.4 percent between 1984 and 2003; in contrast, the Hispanic population more than tripled. If the Hispanic population had grown by only 5.4 percent, but experienced the same increase in the uninsured rate as they actually did over that time (to 40.0 percent uninsured), there would have been 3.1 million uninsured Hispanics in 2003, 6.7 million fewer than the actual figure. Based on our regression estimates, we can only explain 1.8 million (the same figure for both methods), leaving an extra 4.9 million uninsured Hispanics for whom we cannot account from interethnic differences in observable characteristics.

Though our data cannot account for the extra 2.5 to 4.9 million uninsured Hispanics, there are other potential explanations. We present a number of possible factors that could have increased the price of insurance, changed preferences for health insurance, or decreased the income available to purchase insurance for Hispanics during this time.

[28] For Method 1, we multiply the trend coefficient (0.455) times 21 years, so the regression-predicted change in the gap for 2003 is 9.6 percentage points. For Method 2, a change of 9.6 percentage points remains unexplained between 1984 and 2003, as can be seen in the last row of Table 4.

Price of Insurance

Our estimation by source of insurance indicate that most of the unexplained increase in the Hispanic uninsured rate is from fewer Hispanics acquiring coverage through another person. There is also some evidence that employer coverage is harder to come by, though we can explain much of the Hispanic-White difference in employer coverage by controlling for observable characteristics. Perhaps employers are still offering coverage to employees at roughly the same rate, but are attempting to control costs by being more restrictive about coverage for family members. Unfortunately, SIPP does not ask for whether an individual receives an offer of coverage from her employer, and whether that offer will include not just her but also her family members; we are able to determine whether she accepted the offer, but SIPP does not ask whether an offer was rejected.

If jobs that are considered "high-quality" are also the ones who are more likely to offer health insurance for the whole family as part of the compensation package, then perhaps Hispanics have grown more likely to take lower-quality jobs. Whether because of labor market tightness or changing returns to relative skill level, Hispanics could be increasingly likely to be stuck in jobs that offer lower total compensation, less certainty over hours or eventual job tenure, or less workforce bargaining power, all of which would reduce the probability of comprehensive insurance offers. While we find little change in the industrial composition of the Hispanic workforce in our data, it could be that Hispanics find themselves more often in positions that have been less generous with insurance offers within these broad industrial categories.

Also worth considering is the percentage of Hispanics who are self-employed or work in small businesses owned by others, since both groups are likely to face higher and more volatile health insurance plan prices. While SIPP does not have a consistent measure of the size of an employee's firm throughout the years, we do know from our data that the proportion of Hispanics for whom self-employment is their primary job has increased slightly, from 4.53 percent in 1983 to 5.46 in 2003. This is in contrast to the self-employment rate for white non-Hispanics, which grew from 8.5 percent to 9.3 percent in the mid-1990's, but has since fallen below 8.3 percent. Though the share has grown, Hispanics are still much less likely to be self-employed, and this explanation is probably not overly relevant.

It's also possible that Hispanics have increased job volatility. Since many jobs, even "high quality" positions, do not offer health insurance until after a probationary period (often six months or one year), people who change jobs frequently will often experience long spells without health insurance while they wait for eligibility. If, over the time covered in our sample, job tenure declined and job volatility increased for Hispanics, then this could explain part of the growing disparity; we will explore this possibility in future research on insurance transitions.

As with most large survey datasets, we should note that there is no information in the SIPP about an immigrant's legal status. An immigrant lacking proper documentation is likely to find jobs that are of lower quality, as she has little leverage in the job market. Very few of these jobs will offer health insurance for her, and even fewer will extend coverage to her spouse or children. If the proportion of the immigrant population without documentation grew larger from the mid-1980's to the 2000's, then the immigrant population would be less likely to acquire affordable private coverage, in ways that are not captured by our citizenship variables.[29]

Another problem with the SIPP data is the lack of a primary language variable (other than in the 2001 panel). If language or thickness of one's accent is a significant barrier to acquiring health insurance, whether through finding a quality job or discovering eligibility for a public program, many foreign-born (and even some U.S.-born) Hispanics would be more likely to go without coverage. We probably control for most of the language impact by including citizenship variables and their interaction with Hispanic ethnicity and poverty, but within these groups we could still have some with language difficulty having trouble getting insurance, with others with English proficiency being more successful.

For non-English speakers, social networks could prove to be very important when attempting to access health care and health insurance. Gresenz, Rogowski, and Escarce (2007) find that Mexican immigrants are more likely to be insured when they live in a neighborhood with a high concentration of Spanish-speakers and other Hispanics, though this effect disappears for U.S.-born Mexican-Americans, suggesting that those who are more acculturated are less reliant on networks to find access to affordable health care. Also, Borjas and Hilton (1996) find that immigrants are more likely to be covered by Medicaid if earlier immigrants from that same country also had high levels of Medicaid coverage. Our finding of a widening gap in coverage through another person fits into the social network model of these two papers. Still, since network effects are less important for native-born Hispanics, these spillovers are unlikely to explain much of the growing gap with other groups.

Preferences for Risk, Health Care, and Insurance

The lack of insurance among Hispanics could be a matter of personal preference as it relates to risk, in ways that we are not capturing in the variables we have. Barsky et al. (1997) find that in hypothetical questions on a survey, Hispanics are the second-most (after Asians) risk tolerant group, and that

[29] This possible difference in the immigrant population relies on the assumption that SIPP is able to interview the same proportion of illegals that it currently does, and that more illegals will result in more interviews with illegals. On the contrary, if currently no additional undocumented immigrants respond to SIPP interview requests, then there would be no measured difference in the immigrant population.

those with higher risk tolerance are less likely to have health insurance coverage. The survey responses are consistent with increased prevalence of risky behaviors found by other researchers, including lower saving rates (Engen et al. 1999), more credit card usage (Medina and Chau 1998), lower seatbelt usage rates (Campos-Outcalt et al. 2003), and lower vaccination rates (Herrera, Zhao and Klevens 2001).

If native-born Hispanics are healthier, it may make sense for more of them to choose to go without health insurance. But Hendee (1991) finds that Hispanics have a higher risk of diabetes, hypertension, tuberculosis, HIV, alcoholism, cirrhosis, some cancers, and violent deaths. On the other hand, Hispanics may be less likely to seek treatment in an emergency department (Taylor, Larsen, and Correa-de-Araujo 2006), use prescription drugs (Weinick et al. 2004), or have a usual source of care (Weinick, Zurekas, and Cohen 2000), so even if they are less healthy, access to and financial coverage of medical care may be less important. Future research should consider whether Hispanic health care utilization has changed over this time; the decreased likelihood of insurance relative to other groups makes more sense if Hispanics are using less care than they were in the early 1980's, not just that they're less likely to use care at any particular point in time.

Also, Hispanics may value insurance less as a part of their total compensation package, whether it be from an employer or part of a welfare plan. Non-citizens who are remitting income to their families in other countries may prefer to receive compensation in the form of wages rather than fringe benefits, particularly if their time in the U.S. labor market is short or seasonal. This is less likely for citizens, though we should also consider whether Hispanics are more likely to be a secondary wage earner, or have the possibility of being covered by another family member's plan.

Income Available for Insurance

Another possibility is that many Hispanics' lower income makes them unable to afford insurance, even if the price is reasonable and their preference is for coverage. If Hispanics have fallen further behind white non-Hispanics in income over the sample period, then affordability could explain some of the unexplained health insurance disparity. In our data, though, Hispanics have actually gained on white non-Hispanics in family income relative to the federal poverty level. Hispanics' increased likelihood of being in the poorest part of the distribution has shrunk somewhat, and Hispanics are also more likely to be in the 300-400 percent "middle class" range, relative to how the distribution has changed for white non-Hispanics. On the other hand, Hispanics are even less likely than whites to be in the highest income group relative to where they were in the 1980's. Still, the most constrained group seems to have shrunk somewhat, so growing income disparity is likely not the culprit.

Income could be insufficient to purchase insurance if long term or frequent illness cuts into earning power. A person with tuberculosis or cirrhosis would probably be unable to work, or would have frequent interruptions in their work schedule which would make them less attractive to potential employers (not to mention more expensive to insure under a small business employer sponsored plan). The Hendee paper cited above makes the case that Hispanics are more likely to have the type of long term health conditions that could prevent a stable income source.

8 Conclusions

Hispanics were already more likely to be uninsured in the early 1980's. Over the course of the next two decades, they fell even further behind, not only compared to whites, but also relative to blacks, Asians, and Native Americans. Our research suggests that, though differences in education and citizenship between Hispanics and non-Hispanic Whites can account for some of the divergence in health insurance coverage, a substantial portion of the growth in the uninsured rate gap remains unexplained, particularly among individuals of Mexican heritage. In fact, we estimate this portion to represent between 2.5 and 4.9 million extra uninsured Hispanics in 2003. Drilling further down into how health insurance is delivered, we find that most of the unexplained increase was due to lower rates of coverage through a family member, while gaps in public programs and employer-sponsored coverage largely disappear when controlling for observable characteristics.

While we are confident in our findings, we should stress that our paper has some limitations. The biggest downside to using the SIPP is that English language proficiency is not included, except in the 2001 panel. While we expect that we have controlled for most of the effect that language would have on the likelihood of being uninsured with our citizenship status variables, it is likely that English proficiency varies within the Hispanic non-citizen group, and immigrants with better understanding of English will be more likely to find a "quality" job that offers insurance, or better understand their public coverage eligibility.

We are also reliant upon the data being accurate and complete, with only random attrition, though our results are no different when re-weighting observations to control for non-random attrition (see Appendix). Of particular concern is citizenship status; we control for the possibility that those who do not answer the citizenship question could be different than those who do, but it is possible that many non-citizens, especially those without proper documentation, may claim to be citizens, even native-born. We are also unable to tell which immigrants are legal and which are not, which probably has a substantial impact on the likelihood of having health insurance and access to affordable health care.

Our preliminary research indicates that, in addition to the increased likelihood of being uninsured at any particular time relative to other groups, Hispanics have also grown more likely to transition in and out of coverage, though the divergence is not nearly as monotonic as the point-in-time coverage analyzed here. While some medical care can wait until the individual has insurance, frequent transitions may subject policyholders to restrictions on pre-existing conditions, and leave those who have dropped coverage vulnerable to medical and financial risk from poorly timed illnesses or injuries. In the near future, we would like to apply the analysis from this paper to insurance transitions, to determine if the variables we found to be important sources of coverage disparities, such as citizenship and education, also account for the disparity in coverage lapses.

References

[1] Alegria, Margerita, Zhun Cao, Thomas McGuire, and David Takeuchi, 2005. "Health Insurance for Vulnerable Populations: Understanding Differences across Asian American and Latino Subgroups in the United States," Economic Research Initiative on the Uninsured Working Paper 41.

[2] Altonji, Joseph G., and Rebecca M. Blank, 1999. "Race and Gender in the Labor Market," *Handbook of Labor Economics*, Volume 3 (Amsterdam: Elsevier), 3143-3259.

[3] Ashraf, Javed, 1996. "Is Gender Pay Discrimination on the Wane? Evidence from Panel Data, 1968-1989," *Industrial and Labor Relations Review* 49(3): 537-546.

[4] Barsky, Robert B., F. Thomas Juster, Miles S. Kimball, and Matthew D. Shapiro, 1997. "Preference Parameters and Behavioral Heterogeneity: An Experimental Approach in the Health and Retirement Study," *Quarterly Journal of Economics* 112(2): 537-579.

[5] Berk, Marc L., Leigh Ann Albers, Claudia L. Schur, 1996. "The Growth in the U.S. Uninsured Population: Trends in Hispanic Subgroups, 1977 to 1992," *American Journal of Public Health* 86: 572-576.

[6] Blinder, Alan S., 1973. "Wage Discrimination: Reduced Form and Structural Estimates," *Journal of Human Resources* 8(4): 436-455.

[7] Borjas, George J., 2003. "Welfare Reform, Labor Supply, and Health Insurance in the Immigrant Population," *Journal of Health Economics* 22: 933-958.

[8] Borjas, George J., and Lynette Hilton, 1996. "Immigration and the Welfare State: Immigrant Participation in Means-Tested Entitlement Programs," *Quarterly Journal of Economics* 111(2): 575-604.

[9] Bound, John, and Richard B. Freeman, 1992. "What Went Wrong? The Erosion of Relative Earnings and Employment Among Young Black Men in the 1980's," *Quarterly Journal of Economics* 107(1): 201-232.

[10] Buchmueller, Thomas C., Anthony T. LoSasso, Ithai Lurie, and Sarah Dolfin, 2007. "Immigrants and Employer-Sponsored Health Insurance," *Health Services Research* 42(1): 286-310.

[11] Camarota, Steven A., and James R. Edwards., Jr., 2000. "Without Coverage: Immigration's Impact on the Size and Growth of the Population Lacking Health Insurance," Center for Immigration Studies working paper.

[12] Campos-Outcalt, D., C. Bay, A. Dellapena, and M.K. Cota, 2003. "Motor Vehicle Crash Fatalities by Race/Ethnicity in Arizona, 1990-96," *Injury Prevention* 9(3): 251-256.

[13] Carrasquillo, Olveen, David U. Himmelstein, Steffie Woolhandler, and David H. Bor, 1999. "Going Bare: Trends in Health Insurance Coverage, 1989 to 1996," *American Journal of Public Health* 89: 36-42.

[14] Cawley, J. and K.I. Simon, 2005. "Health Insurance Coverage and the Macroeconomy," *Journal of Health Economics* 24(2): 299-315.

[15] Cotton, Jeremiah, 1988. "On the Decomposition of Wage Differentials," *Review of Economics and Statistics* 70: 236-243.

[16] DeNavas-Walt, Carmen, Bernadette D. Proctor, and Jessica Smith, 2007. "Income, Poverty, and Health Insurance Coverage in the United States, 2006," U.S. Census Bureau Report P60-233.

[17] Engen, Eric M., William G. Gale, Cori E. Uccello, Christopher D. Carroll, and David I. Laibson, 1999. "The Adequacy of Household Saving," *Brookings Papers on Economics* 1999(2): 65-187.

[18] Fairlie, Robert W., 2005. "An Extension of the Blinder-Oaxaca Decomposition Technique to Logit and Probit Models," *Journal of Economic and Social Measurement* 30(4): 305-316.

[19] Fairlie, Robert W., and William A. Sundstrom, 1999. "The Emergence, Persistence, and Recent Widening of the Racial Unemployment Gap," *Industrial and Labor Relations Review* 52(2): 252-270.

[20] Fronstin, Paul, Lawrence G. Goldberg, and Philip K. Robins, 1997. "Differences in Private Health Insurance Coverage for Working Male Hispanics," *Inquiry* 34: 171-180.

[21] Gresenz, Carole Roan, Jeanette Rogowski, and Jose J. Escarce, 2007. "Social Networks and Access to Health Care Among Mexican-Americans," NBER Working Paper 13460.

[22] Hamilton, Erin R., Robert A. Hummer, Xiuhong H. You, and Yolanda C. Padilla, 2006. "Health Insurance and Health-Care Utilization of U.S.-Born Mexican-American Children," *Social Science Quarterly* 87(5): 1280-1294.

[23] Hendee, William R., 1991. "Hispanic Health in the United States," *Journal of the American Medical Association* 265(2): 248-252.

[24] Herrera, Guillermo A., Zhen Zhao, and R. Monina Klevens, 2001. "Variation in Vaccination Coverage Among Children of Hispanic Ancestry," *American Journal of Preventive Medicine* 20(4): 69-74.

[25] Jann, Ben, 2005. "Standard Errors for the Blinder-Oaxaca Decomposition," working paper.

[26] Kandilov, Amy M. Gass, 2007. "Health Insurance and Labor Supply Among Recent Immigrants following the 1996 Welfare Reform: Examining the Effect of the Five-Year Residence Requirement," working paper.

[27] Ku, Leighton, and Brian Bruen, 1999. "The Continuing Decline in Medicaid Coverage," *New Federalism: Issues and Options for States*, Urban Institute.

[28] Medina, Jose F., and Chak-Tong Chau, 1998. "Credit Card Usage Behavior Between Anglos and Hispanics," *Hispanic Journal of Behavioral Sciences* 20(4): 429-447.

[29] Monheit, Alan C., and Jessica Primoff Vistnes, 2000. "Race/Ethnicity and Health Insurance Status: 1987 and 1996," *Medical Care Research and Review* 57(Supp. 1): 11-35.

[30] Neumark, David, 1988. "Employers' Discriminatory Behavior and the Estimation of the Wage Discrimination," *Journal of Human Resources* 23: 279-295.

[31] Oaxaca, Ronald, 1973. "Male-Female Wage Differentials in Urban Labor Markets," *International Economic Review* 14(3): 693-709.

[32] Oaxaca, Ronald L., and Michael R. Ransom, 1994. "On Discrimination and the Decomposition of Wage Differentials," *Journal of Econometrics* 61: 5-21.

[33] Reimers, Cordelia, 1983. "Labor Market Discrimination Against Hispanic and Black Men," *Review of Economics and Statistics* 65: 570-579.

[34] Robins, James A., Andrea Rotnitzky, and Lue Ping Zhao, 1995. "Analysis of Semiparameteric Regression Models for Repeated Outcomes in the Presence of Missing Data," *Journal of the American Statistical Association* 90(429), 106-121.

[35] Schur, Claudia, and Jacob Feldman, 2001. "Running in Place: How Job Characteristics, Immigrant Status, and Family Structure Keep Hispanics Uninsured," Commonwealth Fund Working Paper 453.

[36] Shah, N. Sarita, and Olveen Carrasquillo, 2006. "Twelve-Year Trends in Health Insurance Coverage Among Latinos, By Subgroup and Immigration Status," *Health Affairs* 25(6): 1612-1619.

[37] Taylor, Amy K., Sharon Larsen, and Rosaly Correa-de-Araujo, 2006. "Women's Health Care Utilization and Expenditures," *Women's Health Issues* 16(2): 66-79.

[38] Van Hook, Jennifer, Susan L. Brown, and Maxwell Ndigume Kwenda, 2004. "A Decomposition of Trends in Poverty Among Children of Immigrants," *Demography* 41(4): 649-670.

[39] Waidmann, Timothy A., Bowen Garrett, and Jack Hadley, 2004. "Explaining Differences in Employer Sponsored Insurance Coverage by Race, Ethnicity, and Immigrant Status," Economic Research Initiative on the Uninsured Working Paper 42.

[40] Weinick, Robin M., Elizabeth A. Jacobs, Lisa Cacari Stone, Alexander N. Ortega, and Helen Burstin, 2004. "Hispanic Healthcare Disparities: Challenging the Myth of a Monolithic Hispanic Population," *Medical Care* 42(4): 313-320.

[41] Weinick, Robin M., Samuel H. Zuvekas, and Joel W. Cohen, 2000. "Racial and Ethnic Differences in Access to and Use of Health Care Services, 1977 to 1996," *Medical Care Research and Review* 57 (Supp. 1): 36-54.

[42] Wooldridge, Jeffrey M., 2000. "Inverse Probability Weighted M-Estimators for Sample Selection, Attrition, and Stratification," mimeo, Michigan State University Department of Economics.

Appendix

Inverse Probability Weighting for Time Series Graphs

As with most panel surveys, the SIPP experiences nonrandom attrition from the sample. The nonrandom nature of this attrition is a concern for us because evidence suggests that those who are most likely to be uninsured are also the most likely to attrit from the sample. To account for nonrandom attrition, we reweight each observation by the inverse probability of its being present in the sample. Following Robins, Rotnitzky, and Zhao (1995) and Wooldridge (2000), we first determine the probability of individual i being present in the sample at time t using a probit regression, where the covariates are the time $t-1$ (or time-invariant) characteristics and the SIPP interview wave:

$$Prob(present)_{it} = \Phi(Z'_{i,t-1}\gamma),$$

where Φ is the standard normal cumulative distribution function and Z is a vector of personal characteristics. Note that because, by definition, we do not know the value for any time-varying characteristic when the individual is not present, Z includes the value of each characteristic at time $t-1$ (unless, of course, it is time-invariant), with the exception of the SIPP interview wave. In addition to the wave variable, we include dummy variables for being uninsured, unemployed, Hispanic, black, Asian, Native American, female, and married in Z, and use the SIPP weight (lagged as necessary) when running the probit regression.

After fitting a selection probability, π, to each observation, we then find the probability of being present in time t, conditional on being present in each previous period from entry time ($t = 1$) to time $t-1$:

$$p_{it} = \pi_{i1} * \pi_{i2} * \pi_{i3} * ... * \pi_{i,t-1} * \pi_{it}$$

The new bias-corrected weight is then the SIPP-created weight times $1/p_{it}$.

Finally, we rescale all weighted totals so that they match the total number of nonelderly adults in the United States for that month, according to the Census figures.

We use the inverse probability weights in our graphs of the time series of insurance status for different groups, but not in our regressions. Regression results using the weights are not substantially different than the unweighted results, despite SIPP's oversampling of lower income communities. These results are available from the authors upon request. We should also note that we control for the interview wave of the person-month observation in our regressions, in order to pick up some of the potential effect of nonrandom attrition, which is increasingly likely in later waves.

Table A1: Sample Means

	All		Hispanic		White Non-Hispanic	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
Uninsured	0.182	(0.386)	0.367	(0.482)	0.158	(0.365)
Hispanic	0.116	(0.321)	1.000	(0.000)	0.000	(0.000)
Female	0.512	(0.500)	0.523	(0.499)	0.510	(0.500)
Black	0.009	(0.095)	0.079	(0.269)	0.000	(0.000)
Asian	0.001	(0.037)	0.012	(0.108)	0.000	(0.000)
Native American	0.001	(0.038)	0.012	(0.111)	0.000	(0.000)
Married	0.210	(0.407)	0.305	(0.460)	0.198	(0.398)
Own Children Present (0/1)	0.017	(0.130)	0.045	(0.207)	0.014	(0.116)
Age						
18-20	0.078	(0.268)	0.100	(0.300)	0.075	(0.264)

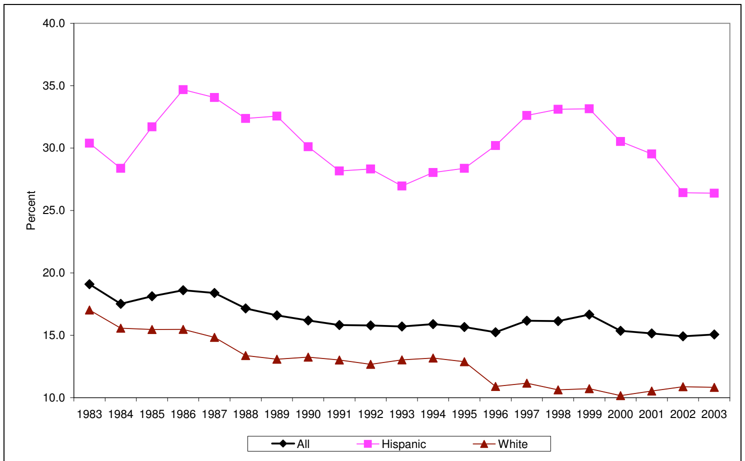
continued on next page

continued from previous page

	All		Hispanic		White Non-Hispanic	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
21-23	0.079	(0.269)	0.096	(0.295)	0.076	(0.266)
24-26	0.083	(0.276)	0.099	(0.298)	0.081	(0.272)
27-30	0.116	(0.321)	0.135	(0.342)	0.114	(0.318)
31-35	0.149	(0.356)	0.159	(0.365)	0.148	(0.355)
36-40	0.148	(0.355)	0.138	(0.345)	0.149	(0.356)
41-45	0.135	(0.342)	0.115	(0.319)	0.137	(0.344)
46-50	0.115	(0.319)	0.090	(0.286)	0.119	(0.323)
51-55	0.097	(0.296)	0.069	(0.254)	0.101	(0.301)
State Unemployment Rate	5.907	(1.711)	6.354	(1.511)	5.848	(1.726)
Citizenship						
N/A	0.181	(0.385)	0.170	(0.376)	0.182	(0.386)
Non-Citizen	0.050	(0.218)	0.311	(0.463)	0.016	(0.125)
Naturalized citizen	0.025	(0.157)	0.098	(0.297)	0.016	(0.124)
Citizenship Interacted with Hispanic						
N/A	0.020	(0.139)	0.170	(0.376)	0.000	(0.000)
Hispanic Non-Citizen	0.036	(0.187)	0.311	(0.463)	0.000	(0.000)
Hispanic Naturalized	0.011	(0.106)	0.098	(0.297)	0.000	(0.000)
Time in U.S.						
Less than 10 Years	0.029	(0.168)	0.172	(0.377)	0.010	(0.101)
10 to 20 Years	0.022	(0.146)	0.129	(0.336)	0.008	(0.086)
More than 20 Years	0.019	(0.137)	0.077	(0.267)	0.012	(0.107)
Less than HS	0.129	(0.336)	0.377	(0.485)	0.097	(0.296)
HS Graduate	0.347	(0.476)	0.315	(0.465)	0.351	(0.477)
Some College	0.235	(0.424)	0.178	(0.383)	0.243	(0.429)
College	0.283	(0.450)	0.122	(0.328)	0.304	(0.460)
Public Coverage Eligibility Proxy						
Under 100% FPL	0.117	(0.322)	0.233	(0.423)	0.102	(0.303)
Female Under 100% FPL	0.068	(0.252)	0.141	(0.348)	0.058	(0.234)
Hispanic Under 100% FPL	0.027	(0.162)	0.233	(0.423)	0.000	(0.000)
Citizenship Interacted with Eligibility Proxy						
N/A	0.029	(0.169)	0.045	(0.207)	0.027	(0.163)
Non-Citizen	0.012	(0.109)	0.086	(0.281)	0.002	(0.047)
Naturalized citizen	0.003	(0.056)	0.017	(0.130)	0.001	(0.036)
Citizenship Interacted with Hispanic and Eligibility Proxy						
N/A	0.005	(0.072)	0.045	(0.207)	0.000	(0.000)
Non-Citizen	0.010	(0.100)	0.086	(0.281)	0.000	(0.000)
Naturalized citizen	0.002	(0.045)	0.017	(0.130)	0.000	(0.000)
Family Income as % of Poverty Line						
Less than 50	0.060	(0.237)	0.100	(0.300)	0.054	(0.227)
50-100	0.058	(0.233)	0.133	(0.340)	0.048	(0.213)
100-150	0.075	(0.263)	0.148	(0.355)	0.065	(0.247)
150-200	0.089	(0.284)	0.138	(0.345)	0.082	(0.275)
200-300	0.187	(0.390)	0.202	(0.402)	0.185	(0.388)

continued on next page

continued from previous page						
	All		Hispanic		White Non-Hispanic	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
300-400	0.164	(0.370)	0.121	(0.326)	0.170	(0.376)
More than 400	0.367	(0.482)	0.157	(0.364)	0.395	(0.489)
Part Time	0.124	(0.329)	0.112	(0.316)	0.125	(0.331)
Full Time	0.582	(0.493)	0.533	(0.499)	0.588	(0.492)
Self Employed	0.085	(0.278)	0.049	(0.217)	0.089	(0.285)
Unemployed	0.609	(0.488)	0.567	(0.496)	0.614	(0.487)
Industry						
N/A	0.224	(0.417)	0.306	(0.461)	0.213	(0.409)
Manufacturing	0.253	(0.435)	0.246	(0.431)	0.254	(0.436)
Utilities	0.050	(0.218)	0.040	(0.195)	0.051	(0.220)
Sales	0.150	(0.358)	0.143	(0.350)	0.152	(0.359)
Service	0.282	(0.450)	0.233	(0.423)	0.289	(0.453)
Government	0.041	(0.197)	0.033	(0.178)	0.042	(0.200)
Spouse Work Status						
Part Time	0.063	(0.244)	0.048	(0.214)	0.066	(0.247)
Full Time	0.355	(0.479)	0.309	(0.462)	0.361	(0.480)
Self Employed	0.064	(0.245)	0.035	(0.185)	0.068	(0.252)
Unemployed	0.119	(0.323)	0.152	(0.359)	0.114	(0.318)
Number of Observations	7,830,614		910,240		6,920,374	
Number of Person Clusters	312,197		38,025		274,172	



Source: Survey of Income and Program Participation, 1984 through 2001 panels; Author calculations.

Figure A1: Uninsured Rate By Ethnicity, Age 17 and Under

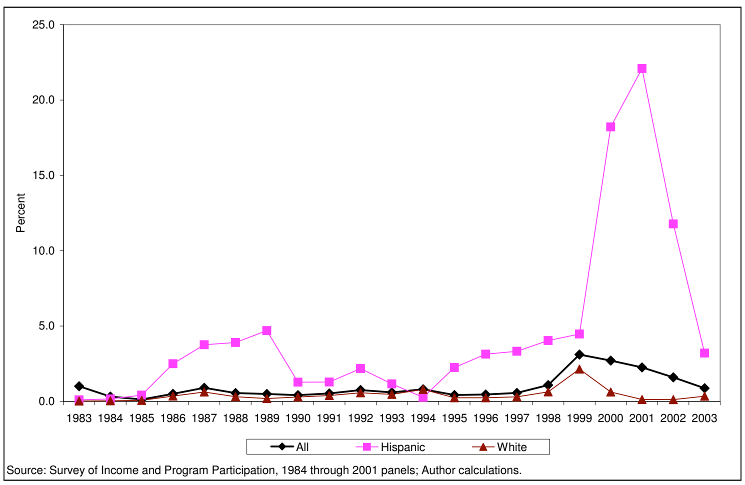


Figure A2: Uninsured Rate By Ethnicity, Age 65 and Over