THE ECONOMIC RESEARCH ORGANIZATION AT THE UNIVERSITY OF HAWAI'I

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Working Paper No. 2019-1

April 4, 2019

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The Impact of Public Health Insurance on Medical Utilization in a Vulnerable Population: Evidence from COFA Migrants

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Abstract

In March of 2015, the State of Hawaii stopped covering the vast majority of migrants from countries belonging to the Compact of Free Association (COFA) in the state Medicaid program. As a result COFA migrants were required to obtain private insurance in health insurance exchanges established under the Affordable Care Act. Using statewide administrative hospital discharge data, we show that Medicaid-funded hospitalizations and emergency room visits declined in this population by 69% and 42% after the expiration of Medicaid eligibility. Utilization funded by private insurance did increase but not enough to offset the declines in publicly-funded utilization. This resulted in a net decrease in utilization. In addition, we show that uninsured ER visits increased as a consequence of the expiration of Medicaid benefits. Paradoxically, we also find a substantial increase in Medicaid-funded ER visits by infants after the expiration of benefits which is consistent with a substitution of ER visits for ambulatory care for the very young.

Key words: Immigration, Health Insurance, Cost Sharing, Medicaid, Insurance Exchange

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JEL Codes: I10, I14, J61

I. Introduction

A principal question in health economics is how insurance coverage affects the demand for health services. Those that lack financial resources are often those most in need of medical services, and, in the absence of adequate insurance, low-income populations may forgo necessary medical care. These concerns have been a driving force for the expansion of government-provided or government-subsidized health insurance in many countries around the world, including the United States. The Affordable Care Act (ACA) of 2010 established subsidies for low-income households to purchase private insurance in marketplaces and incentives for states to expand coverage in their Medicaid programs. As a consequence of the ACA, the percentage of uninsured people in the United States decreased by 41%, a reduction from 48 to 28 million between 2011 and 2015 (Cohen et al, 2017). Importantly, the ACA had the largest impacts on the poor and on minorities (Cohen et al, 2017).

At a time when the United States was expanding insurance coverage for its poorest citizens, the State of Hawaii reduced coverage for a small, but vulnerable, portion of its population. Until March of 2015, the State of Hawaii enrolled eligible migrants of countries belonging to the Compact of Free Association (COFA) in a state Medicaid plan. COFA migrants are from the Republic of Palau, the Republic of the Marshall Islands, and the Federated States of Micronesia (FSM), three nation-states located in the Pacific Ocean.[1] At the time, the State of Hawaii had roughly 28,000 COFA migrants (commonly referred to as "Micronesians"), most of whom were not US citizens, but under the terms of the federal Compacts, are guaranteed certain prerogatives, such as free entry to the US and the right to work. While the Compacts allow for these rights, their allowance for access to health care, particularly Medicaid, has been a highly contested issue.

[1] The Compact of Free Association was signed for the Federated States of Micronesia and the Republic of the Marshall Islands in 1986 (and in 1994 for Palau). Previously, these island groups were under part of the Trust Territory of the Pacific which was administered by the U.S. after World War II. After these countries became independent from the Trust Territory, the citizens of these countries elected to continue their close relationship with the U.S. under a compact of free association (COFA). In exchange for U.S. military access to FSM's ocean territories (an area of over 1 million square miles), the United States agreed to provide governmental funding for the FSM over the course of 15 years; funding was extended beyond that initial time period and is set to expire in 2026. Additionally, FSM citizens were allowed free entry into the United States at any time and, in theory, had access to medical coverage such as Medicaid and other governmental and social services.

In general, Medicaid is jointly financed by federal and state governments. Federal welfare reform in 1996 suspended federal funds for COFA populations through Medicaid. Despite lack of federal Medicaid financing for COFA migrants, the State of Hawaii continued to provide coverage via state-funded health insurance in various forms, including a state Medicaid plan provided by the State of Hawaii Medicaid agency (called Med-QUEST). Due to a court ruling in April of 2014, however, state Medicaid coverage for this population was suspended.[2] As a consequence, most COFA migrants were ultimately denied access to Medicaid benefits in March of 2015. However, some COFA migrants were allowed to maintain coverage. First, CHIPRA (Children's Health Insurance Program Reauthorization Act) continued to cover pregnant women and newborns. Second, the Aged, Blind and Disabled (ABD) program continued to be available to eligible COFA migrants as well. The majority of COFA migrants, however, needed to purchase private insurance in the health insurance exchanges established by the ACA in order to continue their medical coverage; they were not eligible for the Medicaid expansion created by the ACA (McElfish, et al. 2015).

In this paper, we employ statewide administrative data of all hospital discharges in Hawaii to estimate the effects of expiring Medicaid coverage on medical utilization among COFA migrants. The data are close to a census of all hospitalizations in Hawaii over the period 2014-2015. The data also contain a unique patient identification number which enables us to track individual utilization over time. Using these data, we construct an individual-level panel that covers the 24 months from January 2014 to December 2015, which includes months before and after the expiration of Medicaid benefits. The discharge data contain an ethnicity variable. We employ data for three ethnicities: COFA migrants as the treatment group, non-Hispanic whites as the control group, and Japanese as the placebo group. To address omitted zeros for non-utilizers, we include dummy observations and frequency weights corresponding to population numbers obtained from the American Community Survey (ACS).

To investigate the impact of the expiration of Medicaid benefits on utilization among COFA migrants, we use a difference-in-difference research design. We show that there was a sharp reduction in the number of emergency and in-patient medical care admissions charged to Hawaii Medicaid (hereafter referred to as 'Medicaid') after the expiration of benefits for COFA migrants relative to the non-Hispanic white and Japanese populations in Hawaii. In particular, Medicaid-

[2] For details, see McElfish, et al. (2015).

funded ER visits and inpatient admissions declined by 69% and 42%, respectively. This sharp reduction in utilization is consistent with other studies that have investigated the impact of the expiration of Medicaid benefits such as studies on Tennessee after it discontinued Medicaid benefits (see DeLeire 2018, Tarazi 2017, Tello-Trillo 2016). At the same time, there was a substantial increase in the number of emergency room (ER) visits and inpatient admissions charged to private payers, indicating that there was a move towards private insurance among COFA migrants after Medicaid benefits expired. However, the magnitude of this increase was smaller than the reduction in Medicaid-funded utilizations. As a result, net inpatient admissions and emergency visits declined.

Importantly, after Medicaid benefits expired, there was an increase in uninsured ER visits. In particular, for every three COFA migrants who had an ER visit that was paid by private insurance after Medicaid benefits expired, there was a COFA migrant that had an uninsured ER visit. In addition, about one-third of the decline in Medicaid-funded ER visits after the expiration of benefits was made up for by an increase in uninsured ER visits. This is strongly indicative that the effort to enroll COFA migrants in private insurance after the expiration of benefits was not sufficient.

Another key finding of this study is that the expiration of benefits disproportionately impacted infants. We show that inpatient utilization of infants declined dramatically after the expiration of benefits. However, Medicaid-funded ER visits by infants *increased* by a large margin. Due to limitations of the hospital discharge dataset, the precise mechanism underlying this finding is unknown; as noted previously, CHIPRA was extended for this population over this time period. We suspect that one plausible mechanism is that as a consequence of the expiration of benefits, new mothers had fewer out-patient visits for their newborns and, instead, relied on the ER (as opposed to pediatricians) for treating their sick newborns, potentially due to a fundamental misunderstanding of which programs they would continue to be eligible for after the March 2015 reduction in Medicaid coverage. This reduced contact with pediatricians ostensibly would have reduced inpatient admissions since pediatricians typically refer their patients to surgeons and other specialists who work in hospitals.

The main result of this paper is that the expiration of Medicaid benefits reduced utilization on-net for COFA migrants. There are two (related) mechanisms by which this could have happened: increased cost-sharing and low take-up of private insurance. Note that while our study cannot disentangle these mechanisms from each another, the outcome is similar in that patients will consume less medical care when expected out-of-pocket expenses increase.

First, expiration of benefits may have increased the per-unit cost of services since Medicaid has a well-established fee schedule with generally lower reimbursement amounts than private insurance, places restrictions on co-payments, and prohibits balance billing which is the practice of providers charging patients for what insurers do not reimburse. Theoretically, this increase in prices is expected to reduce the consumption of medical services. Indeed, much cited evidence from the RAND Health Insurance Experiment (Manning, et al. 1987; Newhouse, et al. 1993; Aron-Dine et al. 2013) and the Oregon Health Insurance Experiment (Finkelstein et al, 2012) shows that increased cost sharing results in lower utilization. There is also similar quasi-experimental evidence from Card, et al. (2008) in the United States and Shigeoka (2014) in Japan. A reduction in medical utilization as a consequence of an increase in out-of-pocket expenditures without a corresponding underlying change in health status is termed ex post moral hazard or just moral hazard in the health economics literature (Pauly 1968; Cutler and Zeckhauser 2000).

Second, moving COFA migrants from a relatively simple public insurance scheme to more complicated exchanges might have resulted in lower insurance take-up rates (and hence utilization) due to an increase in the complexity of obtaining insurance coverage. In the transition from Medicaid enrollment to private insurance, COFA migrants might have gone uninsured after the expiration of Medicaid benefits, potentially resulting in an increase in out-of-pocket expenses or at least medical charges to individuals that may have gone unpaid. Lower insurance coverage take-up is likely, given that education levels and literacy rates are substantially lower for this population compared to other ethnic groups. For example, Akee (2010) showed that 7.8% of adult male immigrants from the FSM have no education, 6.5% have between one and six years of education, and 16.6% have between seven and eight years of education; the average years of schooling in this population is 10 years. Baicker, et al. (2012) has also showed that take-up rates of low-cost health insurance are low among those of lower income and education levels. Despite considerable outreach by advocates, this take-up issue was widely expected.

As pointed out by Einav and Finkelstein (2018), how medical utilization responds to out-of-pocket costs has been a source of debate for quite some time. Since the RAND Health Insurance Experiment and subsequent studies, the prevailing evidence suggests that the demand for medical care slopes downward. The bulk of our results are consistent with this view. However, a competing view is that the demand for medical care at times slopes *upwards* so that people facing the lowest cost sharing consume the least during the course of year (Einav and Finkelstein, 2018). The idea is that consuming health care that is ostensibly preventive in nature today, can forestall the need for costlier care in the future or that cheaper visits to a primary care physician can substitute for more expensive ER visits.[3] Some ACA-related policy discussions invoked the concept that expanded access to insurance would allow people to consume cheaper preventive care rather seek treatment in the ER.[4] In contrast to this view, however, evidence from the Oregon Health Insurance Experiment (Finkelstein et al, 2012) showed the opposite, namely, that randomized access to Medicaid *increases* ER utilization. We also show that access to Medicaid among COFA migrants increases ER utilization for everyone except for infants. For the youngest COFA migrants, however, there was an increase in ER visits after Medicaid benefits expired, which is a result that is new to the literature.

We conclude that these results provide additional evidence on the responsiveness of the demand for health services to the cost of services in a vulnerable migrant population. The removal of publicly-provided health insurance results in some shifting towards private insurance (possibly under the insurance exchanges established under the ACA). However, this shift does not fully compensate for the decline in utilizations previously financed by Medicaid. In addition, we show that uninsured ER visits have increased as a consequence of the expiration of Medicaid benefits. Overall, our results suggest that there are now COFA migrants forgoing health care services.

The balance of this paper is as follows. In the next section, we provide some institutional background on the history of COFA migrants in Hawaii and their ability to access health insurance. After that, we discuss the discharge data that we employed and how we used it to construct an individual level panel. We then discuss the methods that we employ and after that, we discuss our results and conclude.

II. Some Institutional Background

[3] For example, the model presented in Goldman and Philipson (2007) is very much in this spirit.
[4] See Einav and Finkelstein (2018) for a discussion in the case of Michigan's expansion of its Medicaid program after the implementation of the ACA.

Publicly-sponsored health care coverage for COFA migrants by the State of Hawaii has been subject to various successive federal and state policy decisions. These policy changes resulted in public confusion about the actual health programs and specific benefits for which COFA migrants would be eligible. Such policy changes thus can also serve as a barrier to insurance enrollment and to obtaining health care, further compounding the socioeconomic vulnerability and linguistic and cultural barriers facing this community. Here, we provide a brief overview of recent and relevant policies.

In the 1996 Welfare Reform Act, certain non-U.S. citizens including citizens of COFA nations were deemed ineligible for federal public assistance including Medicaid. Under this Act, immigrants to the US were made ineligible for federal Medicaid assistance unless they have completed a five-year waiting period following immigration in the U.S. However, most COFA residents are not classified as immigrants but instead are classified as legal migrants and specifically permanently classified as "non-qualified aliens." Thus, under this migration status, these individuals are not qualified for federal assistance. To make up for the shortfall in the wake of the 1996 welfare reform, the State of Hawaii began to provide comprehensive health coverage for COFA residents for Medicaid beginning in 1997 using state funds only (Rilkon et al., 2010). However, given that the agreement with COFA nations is a federal not a state policy, the financial responsibility for providing these benefits has often been viewed as disproportionately burdensome to the State of Hawaii, relative to limited federal support available (typically provided through limited Department of Interior funds to different state and territorial jurisdictions affected by COFA migration) (Hawaii DHS, 2009).

After the passage of ACA in 2010, COFA residents along with other lawfully present noncitizens were eligible to purchase health insurance through state health insurance exchanges. However, Medicaid-ineligible noncitizens would not be eligible for federal subsidies for premium-free assistance. Instead, the ACA required Medicaid-ineligible noncitizens with incomes less than 100% of the federal poverty line (FPL) to pay the same premium for insurance purchased on the exchange as a citizen who has income of 100% FPL (Hawaii DHS, 2014).

In the same year, due to budgetary shortfalls the State of Hawaii elected to cancel the Medicaid program eligibility for non-pregnant adult COFA residents. Instead, the State created a limited medical assistance program called Basic Health Hawaii (BHH). Several court cases contesting this change in policy were filed. Following a lawsuit, a federal court issued an injunction “requiring the state to provide Medicaid-like benefits to all non-pregnant adult COFA residents who would otherwise be eligible for Medicaid but for their citizenship status.” The State appealed this injunction to the Ninth Circuit Court of Appeals, which ruled in favor of the State of Hawaii in April 2014. The injunction remained in place until November 2014 when the Supreme Court declined to hear the case, thus ending the plaintiffs’ appeal of the Ninth Circuit decision.

Subsequently, the State created a policy in which non-pregnant adult COFA migrants who were not ABD became ineligible for Medicaid benefits beginning in March of 2015 (Hawaii DHS, 2014; McElfish et al, 2015). Medicaid coverage effectively ended for COFA migrants in the State of Hawaii except for children, pregnant women, and people who were ADB. Infants and pregnant women remained eligible for Medicaid from CHIPRA and those who were ABD were able to receive the same level of benefits as those available under Medicaid.

The non-pregnant, non-ABD COFA adults were told to buy private health insurance on the Hawaii Health Connector, the state’s health insurance exchange. On the exchange, COFA migrants could choose from either of two private insurers (Kaiser Permanente or Hawaii Medical Service Association (HMSA)), with the state paying the premium for insurance purchased for those who were less than 100% of the FPL provided that they chose a Silver-level plan and could verify household income (Hawaii DHS, 2014). The premium assistance program, however, did not pay for any deductible, co-payment, co-insurance, or other cost-sharing arrangements, in contrast to Hawaii Medicaid coverage. However, Kaiser waived these costs for those meeting eligibility requirements by demonstrating financial need.

The final policy shift to insurance exchange plans was a source of much confusion in the community. While outreach volunteers and workers held information sessions and went door-to-door to share relevant information, enrollment on the exchange itself was confusing. Compounding these challenges were the technical challenges troubling the Hawaii Health Connector website, and in 2015, only a few months after the enrollment period, the Connector was closed down and to be replaced by the federally-managed exchange. This meant that anyone who had been enrolled in the connector had to re-enroll using Healthcare.gov, further causing confusion and necessitated outreach to the COFA community (Princeton, 2017). Unlike the Hawaii Health Connector, Healthcare.gov is not available in COFA languages, adding more challenges (Princeton, 2017). In the year 2015 only, the State’s Medicaid program did institute auto-enrollment so those being dropped and those who had not chosen a plan were automatically placed into one of the two private insurance plans, with an intended 50/50 split. A recent policy analysis estimated that 3,600 COFA Hawaii residents enrolled in coverage in Kaiser in 2015 and 5,500 in HMSA (Princeton, 2017).

III. Data Description

The data used in this study are provided by the Hawaii Health Information Corporation (HHIC), a private, not-for-profit organization that was based in Honolulu, Hawaii. HHIC collected data from hospitals in Hawaii. Its catchment area included all hospitals in the State of Hawaii.[5]

We utilized raw data from HHIC that consisted of all utilizations of inpatient and emergency medical services over the period January 1, 2014 to December 31, 2015 for all individuals with Japanese, Caucasian, or Micronesian ethnicities. In total, we used data on 409,556 specific utilizations. For our analysis, we only use utilizations for Hawaii residents (i.e. people with addresses in the State of Hawaii). These data include information on the type of discharge (i.e. inpatient or ER), admission and discharge dates, ethnicity (e.g. Micronesian, Japanese, or Caucasian), gender, age, payer type (e.g. Medicaid, private insurance), total billed/charged, and principal diagnosis and procedural codes. A critical feature of these data is that they include a unique patient identification number which allows us to identify the same patient over time in the raw data. This allows us to construct a panel in which we track utilization of a given individual for each month between January 2014 and December 2015. If no admissions are reported in a given month in the raw data, this indicates that no utilization likely took place in that month given the large catchment area of the HHIC data.

One important feature of the raw HHIC data is that they contain exact birthdates and death dates (for those who died during 2014-15 and provided that they died in a hospital). For people who were born during 2014-15, the panel begins on the month and year of their birth. For people who we know to have died during 2014-15, the panel ends on the month and year of their death.

Descriptive statistics from the raw discharge data are reported in Table 1a. The bulk of the sample is Caucasian comprising 65.6% of all utilizations, followed by Japanese (28.2%) and Micronesians (6.3%). This sample has slightly more women (51.2%) than men (48.8%). Finally,

[5] The HHIC data for Tripler Army Medical Center do not include race information so we do not use hospitalizations and ER visits for this hospital. Accordingly, the data were thus nearly a census.

most of the utilizations in our sample were for people on private plans (32.4%), Medicare (28.3%), and then Med-QUEST (28.0%). Roughly 4% of the utilizations in the raw data were billed to the patient (as opposed to an insurer).

To put the data in a format suitable for regression analysis, we created an individual-level panel in which we tracked utilization for all months between January 2014 and December 2015. To do this, we computed the total number of admissions and charges in a given month for a given individual. We used the discharge date from the raw data to date the utilization. If no utilization took place for an individual in a month, we entered a zero for the cost and utilization variables. Next, we dropped all individual/month observations for which total charge exceeded one million dollars. This resulted in a final panel data set containing 205,691 individuals and 4,782,091 month/individual observations.

The HHIC data and the resulting panel described herein only include individuals with at least one admission to a hospital or an ER during 2014-15. The sample excludes people who had no such contact with the medical system during this time (i.e. people who had no inpatient admission or emergency room visit during this time period). Importantly, if we did observe data for these individuals, the dependent variables (most likely) would have been a 24-month period string of zeros given HHIC's almost universal catchment area.

Table 1b presents population counts from the American Community Survey (ACS). The five-year counts from the ACS from 2011-2015 correspond to people who report Micronesian (excluding Guamanian/Chamorro), Japanese, or White as one of their ethnicities. Our estimates of ACA-based population of Micronesians, Japanese, and Whites are, respectively, 27,890, 310,595, and 604,474, whereas corresponding counts in the HHIC data are 11,530, 63,160, and 131,327 indicating that there are many missing zeros from our panel indicating that they did not use any acute care services.

The solution to this is fairly simple. For each of the three ethnicities considered and for each age/gender category, we added a single dummy observation in which all of the outcome variables were coded as zeros. We then created a set of frequency weights as follows. All individuals in the initial HHIC panel received a weight of unity since they represent exactly one population unit. For each of the dummy observations, which correspond to the omitted zeros from the HHIC data, we set the weight equal to the difference between the population counts for the ethnicity/age/gender category from the ACS and the corresponding ethnicity/age/gender category from the HHIC data. This procedure ensures that the denominators in our means correspond to the population counts as opposed to those who were merely present in the HHIC data (see Appendix A for additional details).[6] Summary statistics on utilization and charges from the panel are reported in Tables 1c and 1d. All statistics use the frequency weights and address the issue of omitted zeros. Table 1c reports statistics for all individuals and Table 1d reports statistics for COFA migrants for the period prior to March 1, 2015. In each of these tables, descriptive statistics are reported for individuals of all ages in the top panel, people under 65 years in the middle panel, and people 65 and over in the bottom panel. Utilization and charges are broken down by inpatient admissions and ER visits. We also report statistics for all utilization under the heading "all payers," utilization charged to Medicaid, utilization charged to private insurance, and utilization not charged to any payer, "uninsured."

In the top panel of Table 1c, we see that on average there were 0.0042 inpatient admissions and 0.0135 emergency room visits per patient-month. This translates to an inpatient admission about every 19 years and an ER admission about every 6 years for entire population. On average, total charges per patient-month for all admissions (i.e. inpatient and ER) were $176.48. The average amount charged to Medicaid was $33.51 and to private payers was $47.48. The remainder was paid by other payers such as Medicare.

Table 1d provides descriptive statistics from the COFA population for the period prior to March 2015. We do this to provide the reader with a baseline to which they should compare the treatment effects that we will compute. The table shows that COFA migrants are sicker than the overall study sample. For example, the mean of hospital admissions per patient-month among COFA migrants under 65 was 0.0075, whereas it was 0.0032 for the entire sample under age 65 in the previous table. Accordingly, the hospitalization rate for COFA migrants is more than twice that of the study sample. Similarly, the lower health status among COFA migrants is also observed by the observation that Micronesians accounted for 6.3% of the study discharges for the three ethnicities, but 3.0% of the state's population of the corresponding ethnicities.

In Figure 1, we display histograms depicting total admissions per patient-month charged to Medicaid, private insurance, Medicare, and the individual by ethnic group for the entire sample period over 2014-2015. The top panel shows inpatient admissions and the bottom panel shows

[6] Note that the counts of many groups in the US Census are not perfect. This is particularly true of marginal groups. Accordingly, we do conduct some sensitivity analysis.

ER visits. The left panel shows utilizations for people under 65 years and the right panel corresponds to people 65 years and older. Several observations are apparent. First, during the duration of our sample, COFA migrants are substantially more likely to have their utilizations charged to Medicaid than either the Japanese or Caucasians. Second, we see a discontinuous jump in total charges to public insurance for people 65 years and older. However, while utilizations of the Japanese and Caucasians are charged to Medicare, COFA migrants are by-and-large covered by Medicaid when they are elderly which is consistent with the discussion in the previous section. Third, COFA migrants visit the ER at much higher rates than the other two groups. Fourth, there is a much higher rate of uninsured COFA migrants in ER usage than either the Japanese or Caucasians (see bottom row of the figure).

Finally, we can use the statistics in Table 1d for COFA migrants to estimate the total cost of providing Medicaid benefits to this population for the State. For all COFA migrants, the State was charged \$234.13 per patient/month for inpatient and ER services. Table 1b indicates that there were 27,890 Micronesians in the State. Accordingly, over the course of a year, this sums to \$78,358,628.40 that was charged to the State for inpatient and ER services. A typical assumption on the payment-charge ratio for Medicaid is 2/3 which results in a total cost of about \$52.5 million to provide inpatient and ER services to COFA migrants. Note that total expenditures by the State of Hawaii in 2014 were about \$10.7 billion (Rosewicz 2018). Accordingly, the cost of providing medical services to COFA migrants for the State constituted approximately 0.5 % of the State's budget in 2014.

IV. Methods

To identify the effects of the expiration of Medicaid benefits on medical utilization among COFA migrants, we employ a difference-in-difference (DD) research design with individual fixed effects. We let y_{it} denote a particular outcome for individual i at month t. The time period in our study is the 24 months between January 2014 and December 2015. The main outcomes that we consider are the total number of inpatient admissions or ER visits in a month and the corresponding total amount charged. We further disaggregate visits by charges to Medicaid, to private insurance, or to the individual.

Our main analysis can be thought of as examining the effect of the change in publicly-provided health insurance on different types of health services. We ask whether there is a change in the total inpatient admissions or ER visits as a result of the program change. We then separate out the outcome variable by whether the subsequent health services were paid by private insurance or Medicaid. Treatment is identified by ethnic group. The treatment group is the COFA population which is identified as “Micronesian” in our data. For convenience, we have chosen Caucasians as the control group and Japanese as placebo group (and the results are the same regardless of designation).

For a given outcome variable y_{it} , the main estimation equation that we employ can then be expressed as:

$$y_{it} = \alpha_i + \vartheta POST_t + \pi POST_t * JP_i + \tau POST_t * COFA_i + g(age_{it}) + \varepsilon_{it} \quad (1)$$

where $POST_t$ is a dummy variable that is equal to one when the calendar month is between March 2015 and December 2015, JP_i is a dummy for an individual of Japanese ethnicity, $COFA_i$ is a dummy that is equal to one if the individual is Micronesian, and age_{it} is the individual’s age at time *t*. The parameter α_i is an individual fixed effect that adjusts for any unobserved time-invariant characteristics that might impact medical demand or might be associated with treatment. The parameter π is the coefficient on the placebo and is expected to be zero. Our parameter of interest is τ, the difference-in-difference estimate of the effect of the change in coverage on medical utilization. We clustered all standard errors by individuals.

We can modify equation (1) to account for a richer form of heterogeneity. Specifically, we consider the following variant:

$$y_{it} = \alpha_i + \rho_i t + \vartheta POST_t + \pi POST_t * JP_i + \tau POST_t * COFA_i + \varepsilon_{it} \quad (2)$$

which includes a heterogeneous time trend given by $\rho_i t$. Note that the heterogeneous trend basically obviates the need for adjusting for age. This specification hedges against some violations of the parallel trends assumption required in the DD model since the trend is allowed to vary across individuals. To estimate this model, we first difference the model to obtain

$$\Delta y_{it} = \rho_i + \vartheta \Delta POST_t + \pi \Delta POST_t * JP_i + \tau \Delta POST_t * COFA_i + \Delta \varepsilon_{it}$$

and, so the coefficient on the trend becomes a fixed effect in the first differenced model. We then apply the standard within group estimator to the first differenced model given above to estimate τ. This specification requires parallel trends in the model in first differences but allows for violations of the parallel trends assumption in levels and, so it is a very robust estimation.

Finally, we also estimate two additional variants of equation (1). The first is a standard event analysis. For these estimations, we include a complete set of time dummies and their interactions with the COFA dummy. Specifically, we estimated the model

$$y_{it} = \alpha_i + \gamma_t + \sum_{s=2}^{24} COFA_i \times \tau_s + g(age_{it}) + \varepsilon_{it} \tag{3}$$

where the s denotes one of 24 months during 2014 and 2015.[7] For each of these estimations, we plot the τ_s estimates for all s. These estimations will also shed light on the parallel trends assumption as the τ_s for the period before March 2015 should be zero if it holds.

The second variant of equation (1) allows us to investigate heterogeneity in the treatment effect by age. For these estimations, we include a complete set of age dummies as well as their interactions with the COFA dummy. For this, we estimated the model

$$y_{it} = \alpha_i + \vartheta POST_t + \sum_{a=1}^{A} (\alpha_a + COFA_i \times \theta_a + COFA_i \times POST_t \times \phi_a) + \varepsilon_{it} \tag{4}$$

where a denotes a five year age bin between ages one and 95. Note that we employed a separate age bin for infants as this is an important age group. This specification features an age profile denoted by the α_a specific to Japanese and Caucasian people, an age profile specific to COFA migrants denoted by θ_a, and the age specific treatment effects denoted by ϕ_a.[8]

V. Results

Core Results

Our core results from the fixed effect DD model are reported in Table 2. The table reports the results of 12 estimations for people under 65 years of age. For each of the estimations, we report the treatment effect on the COFA population as well as the placebo estimate corresponding to the interaction of the post and Japanese dummies. As previously stated, non-Hispanic whites (or Caucasian) is the omitted group for this analysis.

[7] Note that for this specification, we combined the Japanese and Caucasians into the control group. We did this because the subsequent analysis did not indicate that there were different effects on the placebo group. Accordingly, we opted to combine the two groups for the sake of efficiency and parsimony. In addition, the omitted time interaction in the summation is the first period, January 2014.

[8] As before, we are combining Japanese and Caucasian people into a single control group for the sake of parsimony.

Before we proceed, it is important to bear in mind that with the dummy observations, our sample size is effectively about one million individuals which corresponds to the total population of Caucasians, Japanese, and Hawaiians in Hawaii during 2014-2015 (see Table 1b). The vast majority of these individuals are observed for 24 months. Given this large sample size, we would occasionally expect p-values in the vicinity of ten or even five percent even when the true effect of zero due to Type I error (see Deaton 1997). This should be borne in mind when interpreting the estimates on the placebo effects.

In the panel A of Table 2, we look at the effects of the policy change on inpatient admissions and ER visits. For each outcome, we consider admissions funded by any payer, by Medicaid, and by private insurers. In the first column of this panel, we see that the policy had a substantial negative impact on all inpatient admissions with a coefficient estimate of -0.0027 ($p<0.01$) for the COFA migrants and little to no effect in magnitude for the Japanese population, although this estimate is only significant at the 10% level.[9] Note that this effect is inclusive of utilizations that were funded by Medicaid and those that were funded by private insurance. In the fourth column of the same panel, we see a larger reduction in ER visits of 0.0047 ($p<0.01$) per patient-month for the COFA migrants and no effects for the Japanese population. Finally, because we see declines in both inpatient admissions and ER visits together, this is suggestive that the two types of utilizations are complements and are not substitutes, which is consistent with the literature.

Next, looking at utilization disaggregated by type of insurer (columns 2, 3, 5 and 6), we see that utilizations charged to Medicaid declined whereas those charged to private insurance increased for COFA migrants. However, the magnitudes of the former effects are larger than those for the latter effects which is what accounts for the net negative impacts found in the "Any" payer columns (columns 1 and 4). In the second and fifth columns, we see that inpatient admissions and ER visits that were charged to Medicaid declined by 0.0043 ($p<0.01$) and 0.0100 ($p<0.01$) per patient-month. The means of inpatient admissions and ER visits among COFA migrants under 65 that were charged to Medicaid in the pre-policy period were 0.0062 and 0.0240 in Table 1d. Accordingly, these effects amount to 69% and 42% decreases in utilization for the COFA migrants in this time period. In contrast, inpatient admissions and ER visits charged to private insurance increased by 0.0015 ($p<0.01$) and 0.0043 ($p<0.01$), respectively. Compared to the means of 0.0007 and 0.0034 of inpatient admissions and hospital admissions from Table 1d charged to private insurers, these effects represent 214% and 126% increases. These large numbers are entirely attributable to the fact that the vast majority of COFA migrants were not enrolled in private insurance prior to March of 2015. On the whole, this indicates that the policy worked as expected with a shift in financing away from Medicaid and towards private insurers. One important implication is that, since out-of-pocket costs for ER visits tend to be higher for the private insurers (and the uninsured) than for Medicaid, there was a large demand response of the COFA population to higher costs of hospitalizations and ER visits after the implementation of the policy.

The second panel of Table 2 reports the effects for total utilization which is the sum of total inpatient admissions and ER visits in the first three columns as well as total charges (for both ER and inpatient) in the last three columns. The first three columns of the panel (total utilization), mechanically, are the sum of the impacts on inpatient admissions and ER visits from the first panel. Specifically, these are count variables that measure the number of admissions or visits per person. In the first column, we see that, on net, utilization decreased by 0.0074 ($p<0.01$) admissions per patient-month for the COFA migrants. The next two columns indicate that total utilization charged to Medicaid declined by 0.0143 ($p<0.01$) and those charged to private insurers increased by 0.0058 ($p<0.01$). We find no impact, as expected, for the Japanese population.

The final three columns are analogous to the first three columns except that instead of examining the count measure of visits or admissions, we examine the total dollar charges incurred as a result of the expiration of benefits. In the fourth column, we see that on net total (i.e. including both charges to Medicaid and private insurers) charges declined by \$48.48 ($p<0.01$). Breaking these charges down by those charged to Medicaid and to private insurance, we see that charges to Medicaid declined by \$121.21 ($p<0.01$) per patient-month in the second column and those charged to private insurers, in the third column, increased by \$58.04 ($p<0.01$). Using the figures from Table 1d, these correspond a 57% decline and a 239% increase, respectively.

The placebo interactions in Table 2 are generally insignificant. Of the 12 separate estimations, only five of the placebo interactions are significant but mostly only at the 10% level. As argued above, due to our large sample size, this is most likely due to Type I Error. This strongly indicates that our findings are identifying the effects of the expiration of Medicaid benefits on utilization in the COFA population and not an omitted trend.

In the Appendix, we report results from several robustness exercises. First, we estimate the same models as in Table 2 except that we employ alternative weights for COFA migrants. Our preferred weights in this paper are based on a five year count from the ACS from 2011-2015 of Micronesians in the State of Hawaii. The weights that we employ throughout this paper are based off of this count. However, given the possibility that Micronesians (like many other migrant groups) might be undercounted, we also employ the count from the ACS plus the reported margin of error which is 3763. Accordingly, we also report a set of results using weights based off of a count of Micronesians that is slightly higher, 31,653. The results of this exercise are reported in Table A1. Our qualitative findings are unaffected.

The second robustness exercise is reported in Table A2 and uses the log of total admissions and charges as the dependent variable. To account for the large number of zeros in the data, we added one to all observations. Note that because of this, you cannot interpret these estimates as elasticities. We did this to ensure that our findings are not being driven by outliers; this is particularly important for the charges. The results indicate that this is not the case.

The third robustness exercise is reported in Table A3. In this exercise, we compare the estimates of the parsimonious specification from equation (1) that are reported in Table 2 with estimates of a specification with time dummies. The estimates are identical.

In Table 3, we estimate the same models as in the first panel of Table 2 using inpatient admissions and ER visits as the dependent variables with alternative estimators. The first panel extends the fixed effects model estimated in Table 2 to allow for a heterogeneous trend as in equation (2). The second panel employs OLS. Finally, the third panel employs a standard fixed effects model as in Table 2. These estimations are identical to the first panel of Table 2 and are reported for the easier comparisons.

The estimations with the heterogeneous trends in the first panel are broadly consistent with the standard fixed effects estimations in the bottom panel. The key difference is that the standard errors are substantially larger and the point estimates are attenuated. As a consequence, some of the estimates that were significant in the fixed effects model are not significant once we allow for the heterogeneous trend. We attribute this issue to power as the estimates in the first panel of the table essentially rely on a double difference of the original data. That said, the main findings from Table 2 are still present in this more robust model. For example, we see that inpatient admissions charged to private payers increase by 0.0007 ($p<0.10$) per patient-month and ER visits also charged to private payers increase by 0.0028 ($p<0.001$) per patient-month in Panel A for the COFA migrants. Similarly, we see that inpatient admissions charged to Medicaid decline by 0.0013 ($p<0.10$) and ER visits charged to Medicaid decline by 0.0041 ($p<0.05$) per patient-month in Panel A.

In the Panel B of Table 3, we report estimations for the same outcomes using OLS. These estimations tend to be very similar to the fixed effects estimations in the bottom panel of the table with similar (albeit slightly smaller) standard errors.

Next, in Table 4, we estimate the same models as in Table 2 but now we restrict the population to people over 65 years of age. We do not see any impacts of the policy change in this older population. This is the case for both utilizations charged to Medicaid and to private insurers. These results can also be viewed as a placebo test since the policy change did not affect the elderly.

Event Analysis

In Figure 2, we report the results of the event analysis. These figures report the τ_s estimates from equation (3) for all months during the years 2014-2015. The figure contains six graphs corresponding to type of utilization and payer type. The first row displays figures for inpatient admissions and ER visits funded by any payer. The second row displays figures for admissions charged to Medicaid and the final row displays admissions charged to private insurance. All six figures include a horizontal line at zero and a vertical line corresponding to March 2015.

We see the following results. First, the interactions of the month/year dummies with COFA dummy are not significantly different from zero in four of the six graphs prior to March 2015. On the whole, this provides evidence that the trends were parallel across our ethnic groups, at least, in the pre-period. However, the figure in the second row and column corresponding to ER visits charged to Medicaid does show a dip in the pre-treatment period. This result may indicate that the impending change in state coverage of Medicaid for COFA migrants produced a response in emergency room visits prior to the actual change in the policy. Note that the dip that we see for ER visits in the period prior to the expiration of benefits indicates that the estimates in involving ER visits and charges in Table 2 are actual lower bounds (in magnitude) of the true impacts. Second, the first figure corresponding to inpatient admissions charged to any payer shows a small significant decline in the post-period indicating that there was, indeed, a net decline in inpatient utilization as a consequence of the policy. Third, the figure in the second column of the first row which corresponds to ER visits charged to any payer does show a significant decrease in the post-treatment period relative to the pre-treatment period suggesting that there was a net decline in ER visits attributable to the policy change. Fourth, in the second row of Figure 1, we see that both inpatient admissions and ER visits charged to Medicare declined after the expiration of benefits. Finally, the third row indicates utilizations that were charged to private insurers displayed no trends in the pre-treatment period but did increase substantially in the post-treatment period.

Results by Age and Gender

In Figure 3, we investigate how the effects of the expiration of benefits varied by age. In these figures, we report estimates of $\emptyset_a$from equation (4) where the a subscript denotes an age bin. As discussed earlier, the age dummies correspond to five-year age bins, but we did use a separate age bin for children between zero and one year of age. For these estimations, we include individuals both under and over age 65. Each figure contains a horizontal line at zero and a vertical line at age 65 after which the policy change should have had no impact.

The first row displays results for inpatient admissions and ER visits charged to any payer. The first figure corresponds to inpatient admissions and shows a flat profile close to zero with the exception of infants for whom we see a large negative effect. These age effects are very tightly estimated until 65 years after which the confidence bands around them become substantially larger. The second figure in the row shows a net decrease in ER visits for most ages up until age 65 at which point the age effects are not significantly different from zero. The bulk of the decline in ER visits occurs for people between the ages of 30 and 65. However and importantly, we see an *increase* in ER visits for Micronesian infants *after* the expiration of Medicaid benefits; this is indicated by the observation at one year of age (which corresponds to infants) which is positive and statistically significant.

In the second row of Figure 3, we display results for inpatient admissions and ER visits charged to Medicaid. The two figures show that inpatient admissions of infants declined and ER visits of infants increased. Both did so in a dramatic fashion. Note that the negative effects on infants' inpatient admissions are very large relative to the negative effects on other COFA migrants under 65 years. In this figure it is difficult to detect, but the effects for COFA migrants between age one and 65 are in fact significantly different from zero in many instances. This is more apparent in Figure A1 in which we plot the same estimates from the first figure in the second row with the omission of the infants.

The contrasting effects of the expiration benefits on the utilization of infants are important. We showed that inpatient admissions declined precipitously whereas ER visits increased by a large magnitude. While our hospital data is perhaps not ideal for pinning down the precise mechanism, we suspect that the expiration of Medicaid benefits led to a decline in the use of ambulatory care by new mothers which probably led to two consequences. The first is that the use of ER visits for neonatal care increased. The second was a decline in inpatient admissions, which may have happened since primary care physicians refer patients for surgery and other inpatient services. *Prima facie*, these effects on infants are puzzling since pregnant women and legally residing children were still technically covered by Medicaid after March, 2015 (i.e. the policy change should not have affected this group). What these results suggest then is that despite their continued eligibility, there were still many children and pregnant mothers that did not did not continue to use the available state-provided Medicaid services – perhaps because they were unaware that they continued to be eligible for CHIPRA coverage.

In some sense, this can be viewed as a reverse woodworking effect. Benefits expired for a large swath of the Micronesian population in Hawaii. Many COFA migrants were actually still covered by the State's Children's Health Insurance Program (CHIP) program. However, it appears as if the salience of the expiration of benefits for the majority of migrants led many eligible migrants to believe that they were not covered. In a similar vein, but in the opposite direction, Frean, et al. (2017) found that the expansion of Medicaid under the ACA increased enrollment in Medicaid among people who were previously eligible for Medicaid benefits.

The final row of Figure 2 displays the results for utilization charged to private payers. Both figures in this row show modest increases in both inpatient admissions and ER visits for individuals below age 65. Importantly, the magnitude of the effects for ER visits charged to private insurers are substantially smaller than the corresponding effects for admissions charged to Medicaid indicating that the policy change did reduce utilization of emergency medical services. Finally, with private insurers, we also see that there was a substantial decline in inpatient admissions and a substantial increase in ER visits.

In Table 5, we examine the impacts across gender. We report DD estimates of the effects of the policy on inpatient and ER visits charged to any payer, Medicaid, and private insurance. We further restrict the estimations to people under 65. The top panel displays impacts on females and the bottom panel displays impacts on males.

The main result in this table is that COFA females were affected more than COFA males. For example, female COFA migrants saw a decline in inpatient admissions charged to Medicaid of 0.0047 ($p<0.01$) per patient-month, whereas males saw a decline of 0.0040 ($p<0.01$) per patient-month. The corresponding estimates for ER visits charged to Medicaid are -0.0126 ($p<0.01$) for females and -0.0072 ($p<0.01$) for males which again indicates an impact on females that is about twice as large as it is for males and these two estimated coefficients are statistically significantly different from one another.

Effects on the Uninsured

We now look at how the uninsured were impacted by the policy change. Before we proceed, it is important to note that it was widely suspected that many migrants were not able to enroll in and/or maintain coverage in private insurance following this policy change. To shed light on this, we estimate our fixed effects DD model using inpatient admissions and ER visits that were not charged to any insurer (either public or private) as the dependent variable. We report the results in Table 6. The table displays four DD estimates corresponding to inpatient and ER visits for people who are under 65 years and separately for those who are 65 years and older.

We see some impacts for COFA migrants under 65 years, but not older than 65. There were no effects on uninsured inpatient admissions. However, we do see that uninsured ER visits increased by 0.0016 ($p<0.01$) per patient-month. Note that the impact on uninsured ER visits is one third of the impact on ER visits charged to private insurers of 0.0043 from Table 2. This is an important result as it indicates that efforts after the expiration of Medicaid benefits to enroll people were private insurance were not effective. For every three ER visits by COFA migrants that were paid by private insurance (after Medicaid benefits expired), there was an ER visit by a COFA migrant that was not paid by insurance. In addition, the impact on ER visits charged to Medicaid in the same table was -0.0043. Accordingly, the reduction in insured COFA migrants'

ER visits was offset by an increase in the number of uninsured visits to ER by COFA migrants; the corresponding increase in uninsured ER visits is about one-third of the decline in Medicaid funded ER visits. The last two columns show no impacts on uninsured admissions among older COFA migrants.

In Figures 4a and 4b, we report similar event analyses and age profiles of the DD estimates for uninsured admissions as we did in Figures 2 and 3. First, in the event analysis in Figure 4a, we see that we very large impacts of the policy on uninsured ER visits. However, a large portion of this increase pre-dates March 2015. It is not completely clear what the cause of this ramp-up is for pre-March 2015. In keeping with this story, it is important to note that the plot of Medicaid-charged ER visits from Figure 2 also shows a decline prior to the official expiration date. This suggests that prior to the "official" expiration date of March 2015, there was a substitution away from Medicaid-charged ER visits towards uninsured ER visits. Typical practice at most hospitals with emergency rooms is to enroll eligible, uninsured patients in Medicaid; this guarantees that the hospital will get paid for the visit. That we do not see this in the period just prior to the expiration date indicates a rigidity preventing this from happening.[10] Second, in Figure 4b, we see that the bulk of the impacts on uninsured ER visits occurred for adults younger than 65. Note that we do not see large impacts on uninsured ER visits for infants, nor were there impacts on COFA migrants 65 years and older. Finally, the effects on uninsured hospital admissions were substantially smaller.

Effects on the Infants

Finally, we further explore how the expiration of benefits impacted Micronesian infants. To do this, we estimate a variant of equation (1) in which we include an infant dummy and its interaction with the COFA/POST variable. The idea of this is to shed light on how the policy impacted infants in a transparent way. We report the results in Table 7. We only consider admissions that were charged to Medicaid. The table consists of eight columns corresponding to four outcomes: inpatient admissions, ER visits, and inpatient charges and ER charges. For each outcome, we report the estimation of equation (1) which was also reported in Table 2 in the odd columns and, in the even columns, we report the results with the infant interaction.

[10] Private communication with physicians working at Queens Medical Center in Honolulu indicated that just prior to the expiration of Medicaid benefits, there was a sense that it would be difficult to enroll uninsured COFA migrants in the State's Medicaid program so many providers may not have put forth the effort.

The results indicate that the effects of the Medicaid expiration on infants were enormous. Looking at the first two columns where the dependent variable was inpatient admissions paid by Medicaid, we see that the raw impact of the policy was -0.0043 ($p<0.01$) in the first column. In the second column, the direct impact declines to -0.0033 ($p<0.01$), but the interaction with the infant dummy is -0.0978 ($p<0.01$), so that effect of the policy change was substantially larger for infants. Accordingly, the net effect on infants was -0.0978-0.0033=-0.1011 inpatient admissions per patient-month. We see a similar phenomenon in the third and fourth columns where ER visits are the dependent variable in that infants were more heavily impacted than the rest of the population. However, with ER visits, we see that the expiration of benefits caused Medicaid-funded ER visits to increase by a large margin. The fourth column shows a direct effect of the policy on ER visits of -0.0106 ($p<0.01$), but the interaction with the infant dummy is 0.0617 ($p<0.01$). Hence, the net increase on ER visits for infants is 0.0617-0.0106 = 0.0511 ER visits per patient/month. The final four columns of the table report the effects on charges. These results indicate that the policy decreased charges for inpatient admissions of Micronesian children by -1147.46-87.42 = -1234.88 dollars per patient-month. The corresponding number for ER visits was 86.92 -22.61 = 64.31 dollars per person/month.

VI. Conclusions

In this paper, we investigated the effects of eliminating Medicaid coverage for a vulnerable migrant population in the State of Hawaii. To do this, we employed a large administrative database that constitutes close to a census of all inpatient and emergency room utilizations during 2014 and 2015. Difference-in-difference models indicate that the expiration of benefits decreased Medicaid-funded inpatient and emergency room utilizations by 69% and 42%, respectively. Privately-funded utilizations increased by 214% for inpatient admissions and 126% for emergency room visits. On net, the magnitudes of the publicly-funded utilization did not make up for the decline in Medicaid-funded utilization resulting in a net decline in utilization after the expiration of Medicaid benefits.

Some of the shortfall in Medicaid-funded utilization of the ER was made up for by utilizations of uninsured patients. We find that there was a marked increase in ER visits that were charged to the patient (as opposed to Medicaid) that began earlier than the official expiration date of Medicaid benefits. This is a puzzle. Ostensibly, COFA migrants should have been eligible for Medicaid benefits up to March of 2015. Our best guess is that COFA migrants who were enrolled in Medicaid were allowed to obtain Medicaid benefits until March of 2015. However, we also suspect that the perception that Medicaid benefits were ending resulted in the dramatic reduction in enrollment prior to March 2015; this may have been driven by the patients or the providers. Effectively, this meant that Medicaid benefits expired for uninsured COFA migrants prior to their official expiration date.

Despite our finding that Medicaid-funded utilization declined, we did see that there was a dramatic *increase* in Medicaid-funded ER visits by Micronesian infants *after* the general Medicaid benefits expired. Unfortunately, it is hard to pin down the precise mechanism underlying this finding, but we suspect that Micronesian parents substituted ER visits for ambulatory care for their newborns once Medicaid benefits expired. However, what is puzzling about this is that Micronesian newborns were (and are) still eligible for Medicaid via the State's CHIPRA program. If ER visits were in fact substituting for ambulatory care for Micronesian infants, this is suggestive of a failure to effectively communicate that the children of COFA migrants would continue to be eligible for Medicaid even after benefits for most other COFA migrants ended.

Many of the undesirable effects of the expiration of benefits were predicted at the time of expiration. For example, Hagiwara, et al. in the May 2015 issue of the *Journal of Health Care for the Poor and Underserved* said, "There is concern that this process, which has proven to be confusing even for native English speakers, will at best be confusing for COFA migrants and at worst cause individuals to be uninsured and possibly forgo needed health care." These prognostications turned out to be true.

An important take-away of this study for policy makers is that moving poorer people from Medicaid to private insurance obtained from exchanges may result in lower utilization on net. This could happen for three reasons. First, the relative complexity of the exchanges could result in lower take-up rates of private insurance thereby leaving many without insurance. Second, the vast majority of private insurance plans entail more out-of-pocket expenses than Medicaid which typically has little or no out-of-pocket expense. Third, communicating these changes and options is not a trivial undertaking. For vulnerable populations with limited English ability and familiarity with government agencies, this may prove to be a larger hurdle than for the native-born populations. For these reasons, we would expect a transition from Medicaid to private insurance to reduce medical demand or utilization.

In the context of the subject of this study, better communications could have improved the ease of the transition from public to private insurance in three ways. First, the State of Hawaii could have better communicated when the enrollment period for Medicaid ended for the majority of COFA migrants – March 2015. The ramp-up in uninsured people visiting the ER prior to the official March 2015 expiration of Medicaid benefits indicates an institutional impediment preventing enrollment into Medicaid during a period in which COFA migrants ostensibly were eligible for Medicaid benefits. A better way to have implemented the transition would have been to announce a date in which enrollment into Medicaid would end as opposed to a date at which benefits would expire. After the enrollment period ended, the State should have commenced a more aggressive program to enroll COFA migrants into private insurance via the exchanges. Second, the State should have better communicated who was eligible for Medicaid benefits after they expired for the bulk of COFA migrants. Our findings are strongly consistent with a substitution of ER visits for ambulatory care for Micronesian infants after Medicaid benefits expired. It is likely that better communications could have educated COFA migrants that they could still obtain publicly-funded ambulatory care for their children after March of 2015. Finally, the State could have better communicated how to properly navigate the exchanges to obtain private insurance coverage. The sharp decline in utilization after the expiration of Medicaid benefits was due to either cost-sharing effects and/or low take-up rates of private insurance. Employing better communications to improve take-up of private insurance most likely would have mitigated the observed decline in utilization that we estimated in this study. It also would have likely mitigated the increase in uninsured ER visits that occurred after the benefits expired.

While this study focuses on a very unique policy change affecting a relatively small population, we claim that we can provide lessons to other policy makers. In particular, we have shown the difficulties of using private insurance obtained through exchanges to provide coverage to vulnerable migrant populations with low levels of education and English proficiency. On the whole, we suspect that a relatively simpler single payer public insurance scheme would be better suited for such a population. However, if policy makers are insistent on using private insurers to cover vulnerable migrant populations, better communications and outreach are needed. Hopefully, this study provides some lessons on how to proceed.

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Appendix A: Discussion of Weighting Procedure

We let $S_i \in \{0,1\}$ denote an indicator for being present in the HHIC data where unity indicates presence. Focusing on the basic DD model in equation (1) and abstracting from the age function and the Japanese placebo group for the ease of notation, the DD parameter is

$$\tau = \{E[y_{it}|POST_t = 1, COFA_i = 1] - E[y_{it}|POST_t = 0, COFA_i = 1]\}$$
$$- \{E[y_{it}|POST_t = 1, COFA_i = 0] - E[y_{it}|POST_t = 0, COFA_i = 0]\}.$$

However, given our sampling issues, without any adjustment, in the HHIC data, we can identify the parameter

$$\tau^* = \underbrace{\{E[y_{it}|POST_t = 1, COFA_i = 1, S_i = 1] - E[y_{it}|POST_t = 0, COFA_i = 1, S_i = 1]\}}_{\Delta_1}$$
$$- \underbrace{\{E[y_{it}|POST_t = 1, COFA_i = 0, S_i = 1] - E[y_{it}|POST_t = 0, COFA_i = 0, S_i = 1]\}}_{\Delta_0}.$$

which, in general is not equal to τ. However, it turns out that

$$\tau = \omega_1\Delta_1 - \omega_0\Delta_0$$

where $\omega_d \equiv \frac{P(COFA_i=d, S_i=1)}{P(COFA_i=d)}$. For $d = 1$, the numerator in the weight is the probability of being a COFA migrant and in the HHIC data. The denominator is the probability of being a COFA migrant.

To see this, note that we can write

$$E[y_{it}|POST_t = p, COFA_i = d, S_i = 1] = \frac{E[y_{it} \times 1(POST_t = p, COFA_i = d, S_i = 1)]}{p(POST_t = p, COFA_i = d, S_i = 1)}$$

for $p, d \in \{0,1\}$. Next, we note that

$$E[y_{it} \times 1(POST_t = p, COFA_i = d, S_i = 1)] = E[y_{it} \times 1(POST_t = p, COFA_i = d)]. \quad \text{(1a)}$$

This is true because $S_i = 0$ implies that $y_{it} = 0$. Accordingly, we obtain that

$$\frac{p(POST_t = p, COFA_i = d, S_i = 1)}{p(POST_t = p, COFA_i = d)} E[y_{it}|POST_t = p, COFA_i = d, S_i = 1] \tag{1b}$$
$$= E[y_{it}|POST_t = p, COFA_i = d]$$

.

Finally, if we assume that the likelihood of being in the COFA population is the same in the pre- and post-periods, then we obtain that

$$\omega_d E[y_{it}|POST_t = p, COFA_i = d, S_i = 1] = E[y_{it}|POST_t = p, COFA_i = d]$$

where $\omega_d \equiv \frac{P(COFA_i = d, S_i = 1)}{P(COFA_i = d)}$.

An important insight from the calculations above (namely equations (1a) and (1b)) is that

$$E[y_{it}|POST_t = p, COFA_i = d] = \frac{E[y_{it} \times 1(POST_t = p, COFA_i = d, S_i = 1)]}{p(POST_t = p, COFA_i = d)}.$$

This is interesting because it suggests that the expectation on the left-hand side of the above equation can be estimated as

$$\sum_{i=1}^{N} y_{it}\varphi_i$$

where $\varphi_i = \frac{f_i}{\sum_{i=1}^{N} f_i}$ and f_i is the frequency weight associated with the ith observation. Note that N corresponds to the sample size in the HHIC data with the added dummy observations discussed in Section 3. The frequencies are equal to unity if the observation in in the HHIC data. For the dummy observations, they are equal to the difference between the counts from the ACS and the HHIC sample for a given ethnicity/age/gender cell. An important feature of the weights in φ_i is that they exactly correspond to what standard statistical packages such as STATA compute when you employ frequency weights.

The frequency weights that we employ are based off of the five-year population counts from the ACS reported in Table 1b. We used these counts to construct frequencies for gender/age/period/ethnicity cells. Note that the sample of Micronesians in Hawaii in any given year of the ACS is quite small, but five year averages of the ACS can be used to arrive at a fairly reliable aggregate population count. However, as pointed out by Fernandez, et al. (2018), using the ACS to construct precise counts of specific age groups and, particularly, the very young is very difficult. This is especially true for relatively small groups such as Micronesians as there is only about 200-300 Micronesians in Hawaii total in any given ACS year. Accordingly, we used the ACS to compute the proportions of Japanese, Caucasian, and Micronesian females and males who were under and over 65 which is a relatively broad category. We then used these proportions to count the numbers of each ethnicity/gender category under and over 65 and took the difference between these counts and the counts in the HHIC data. These numbers constitute the number of omitted zeros in each gender/ethnicity category under and over age 65. The resulting number of omitted zeros was then evenly allocated to each age between zero and 85.

Table 1a: Descriptive Statistics for the Raw Data

	Counts [Percentage]
Race	
Japanese	115,456 [28.2%]
Micronesian	25,621 [6.3%]
White	268,479 [65.6%]
Gender	
Male	199,758 [48.8%]
Female	209,796 [51.2%]
Unknown	2 [0.0%]
Payer Type	
Department of Defense	19,132 [4.7%]
Medicaid/Quest	114,711 [28.0%]
Medicare	115,907 [28.3%]
Miscellaneous	9,886 [2.4%]
Private Insurance	132,601 [32.4%]
Self-pay	17,319 [4.2%]

Notes: These are tabulations from the raw discharge data that we used to construct the final panel. The raw data consisted of 409,556 utilizations.

Table 1b: Population Counts in the HHIC and ACS Data

Self-reported ethnic group	HHIC	ACS	HHIC/ACS
Micronesian*	11,530	27,890	41%
Japanese	63,160	310,595	20%
White/Caucasian	131,327	604,474	22%

Notes: We used the American Community Survey over the years 2011-2015 to compute the population numbers for a given year. The counts from the ACS account for people reporting multiple races.
*Excludes Guamanian/Chamorro.

Table 1c: Descriptive Statistics from Panel Data

	Any Payer	Medicaid	Private	Uninsured
		All		
Inpt. Admissions	0.0042	0.0009	0.0013	0.0001
	(0.0683)	(0.0309)	(0.0380)	(0.0081)
ER Visits	0.0135	0.0041	0.0045	0.0007
	(0.1384)	(0.0817)	(0.0733)	(0.0291)
Total Charges	176.48	33.51	47.48	2.89
	(4032.16)	(1802.81)	(2006.49)	(309.38)
Inpt. Charges	144.72	25.23	37.05	1.39
	(3989.21)	(1777.83)	(1985.40)	(293.75)
ER Charges	31.76	8.27	10.44	1.50
	(411.12)	(212.28)	(217.65)	(88.54)
		Under 65 years		
Inpt. Admissions	0.0032	0.0011	0.0016	0.0001
	(0.0593)	(0.0347)	(0.0415)	(0.0090)
ER Visits	0.0138	0.0053	0.0055	0.0009
	(0.1405)	(0.0924)	(0.0813)	(0.0327)
Total Charges	121.55	41.43	54.68	3.47
	(3313.44)	(1988.17)	(2124.44)	(325.91)
Inpt. Charges	91.53	30.85	42.02	1.59
	(3269.57)	(1959.67)	(2100.98)	(307.55)
ER Charges	30.03	10.58	12.66	1.88
	(394.10)	(239.40)	(238.56)	(98.52)
		65 years and older		
Inpt. Admissions	0.0076	0.0001	0.0005	0.00002
	(0.0916)	(0.0120)	(0.0227)	(0.0045)
ER Visits	(0.0126)	0.0003	0.0011	0.00009
	(0.1313)	(0.0235)	(0.0369)	(0.0107)
Total Charges	354.87	7.78	24.10	1.04
	(5776.08)	(983.43)	(1562.97)	(248.22)
Inpt. Charges	317.47	6.98	20.89	0.76
	(5730.63)	(976.22)	(1551.64)	(243.54)
ER Charges	37.40	0.80	3.22	0.28
	(462.08)	(72.20)	(127.80)	(42.18)

Notes: Reports means and standard deviations in parentheses. All statistics are on a per patient/month basis. Dummy observations and frequency weights were used to account for missing zeros.

Table 1d: Descriptive Statistics from Panel Data: COFA Migrants in Pre-Treatment Period

	Any Payer	Medicaid	Private	Uninsured
		All ages		
Inpt. Admissions	0.0078	0.0065	0.0007	0.0004
	(0.0911)	(0.0832)	(0.0263)	(0.0214)
ER Visits	0.0322	0.0238	0.0032	0.0044
	(0.1951)	(0.1685)	(0.0611)	(0.0709)
Total Charges	289.42	234.13	24.32	19.77
	(5331.29)	(5049.35)	(1217.52)	(747.32)
Inpt. Charges	226.42	190.12	17.01	10.32
	(5287.91)	(5016.62)	(1196.79)	(709.79)
ER Charges	63.00	44.01	7.31	9.45
	(513.87)	(412.66)	(184.76)	(214.65)
		Under 65 years		
Inpt. Admissions	0.0075	0.0062	0.0007	0.0005
	(0.0891)	(0.0813)	(0.0266)	(0.0216)
ER Visits	0.0327	0.0240	0.0034	0.0046
	(0.1964)	(0.1694)	(0.0621)	(0.0722)
Total Charges	263.92	210.89	24.31	19.85
	(5057.99)	(4772.65)	(1222.47)	(735.73)
Inpt. Charges	201.74	167.84	16.90	10.15
	(5014.97)	(4740.65)	(1203.42)	(696.47)
ER Charges	62.19	43.05	7.41	9.70
	(502.42)	(400.56)	(179.24)	(217.08)
		65 years and older		
Inpt. Admissions	0.0135	0.0114	0.0005	0.0004
	(0.1221)	(0.1126)	(0.0212)	(0.0188)
ER Visits	0.0238	0.0189	0.0012	0.0016
	(0.1683)	(0.1504)	(0.0400)	(0.0407)
Total Charges	747.14	651.50	24.54	18.29
	(8898.39)	(8608.45)	(1124.95)	(931.26)
Inpt. Charges	669.54	590.16	19.04	13.38
	(8845.47)	(8563.99)	(1071.00)	(916.59)
ER Charges	77.60	61.34	5.50	4.90
	(687.62)	(588.75)	(264.92)	(165.02)

Notes: Reports means and standard deviations in parentheses. All descriptive statistics correspond to the period prior to March 1, 2015. All statistics are on a per patient/month basis. Dummy observations and frequency weights were used to account for missing zeros.

Table 2: Fixed Effects DD Estimates: Under 65

Panel A	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0002*	0.0001	-0.0003**	-0.0002	-0.0002*	0.0002
Population	(0.0001)	(0.0000)	(0.0001)	(0.0002)	(0.0001)	(0.0001)
COFA	-0.0027***	-0.0043***	0.0015***	-0.0047***	-0.0100***	0.0043***
Population	(0.0003)	(0.0003)	(0.0001)	(0.0006)	(0.0007)	(0.0003)

Panel B	Total Utilization			Total Charges		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0004*	-0.0001	-0.0001	-9.8973*	-1.8706	-4.6185
Population	(0.0002)	(0.0001)	(0.0001)	(4.2818)	(2.0704)	(3.2191)
COFA	-0.0074***	-0.0143***	0.0058***	-48.4754***	-121.2113***	58.0373***
Population	(0.0007)	(0.0009)	(0.0004)	(14.4764)	(13.6581)	(6.9071)

*significant at the 10% level; **significant at the 5% level; ***significant at the 1% level

Notes: All estimations use 156,944 individuals (including the dummy observations) observed over a maximum of 24 months and include individual fixed effects, a dummy for the post-policy period, and a quadratic function of age as controls. Japanese Population is the interaction of the post dummy with the Japanese dummy. COFA Population is the interaction of the post dummy with the COFA dummy. Non-Hispanic White (Caucasian) is the omitted reference group in this analysis. Standard errors adjust for clustering on individuals.

Table 3: Alternative Estimations: Under 65

Panel A First Differenced Fixed-Effects

	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	0.0000	0.0000	0.0001	-0.0005	-0.0003	-0.0001
Population	(0.0002)	(0.0001)	(0.0002)	(0.0005)	(0.0002)	(0.0003)
COFA	-0.0003	-0.0013*	0.0007*	0.0002	-0.0041**	0.0028***
Population	(0.0007)	(0.0006)	(0.0003)	(0.0018)	(0.0014)	(0.0007)

Panel B OLS

	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0001	-0.0000	0.0000	-0.0002	-0.0002	0.0002
Population	(0.0001)	(0.0000)	(0.0001)	(0.0002)	(0.0001)	(0.0001)
COFA	-0.0009**	-0.0024***	0.0015***	-0.0027***	-0.0080***	0.0042***
Population	(0.0003)	(0.0002)	(0.0001)	(0.0006)	(0.0006)	(0.0003)

Panel C Fixed-Effects

	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0002*	0.0001	-0.0003**	-0.0002	-0.0002*	0.0002
Population	(0.0001)	(0.0000)	(0.0001)	(0.0002)	(0.0001)	(0.0001)
COFA	-0.0027***	-0.0043***	0.0015***	-0.0047***	-0.0100***	0.0043***
Population	(0.0003)	(0.0003)	(0.0001)	(0.0006)	(0.0007)	(0.0003)

*significant at the 10% level; **significant at the 5% level; ***significant at the 1% level
Notes: Per Table 2.

Table 4: Fixed Effects DD Estimates: Over 65

Panel A	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0000	0.0000	0.0001	0.0001	0.0000	0.0003**
Population	(0.0003)	(0.0000)	(0.0000)	(0.0003)	(0.0000)	(0.0001)
COFA	-0.0010	-0.0014	0.0003	-0.0006	0.0002	0.0003
Population	(0.0015)	(0.0013)	(0.0003)	(0.0023)	(0.0021)	(0.0004)

Panel B	Total Utilization			Total Charges		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	0.0001	0.0000	0.0003**	-1.8561	-0.0677	0.3412
Population	(0.0005)	(0.0000)	(0.0001)	(14.4428)	(1.3169)	(3.0298)
COFA	-0.0016	-0.0012	0.0005	-118.3847	-128.4211	-0.4587
Population	(0.0030)	(0.0027)	(0.0005)	(101.3261)	(96.3382)	(10.9350)

*significant at the 10% level; **significant at the 5% level; ***significant at the 1% level

Notes: All estimations use 54,398 individuals (including the dummy observations). All other notes per Table 2.

Table 5: Fixed Effects DD Estimates by Gender: Under 65

Panel A: Females

	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0001	0.0001*	-0.0001	-0.0004	-0.0004**	0.0004*
Population	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0002)	(0.0002)
COFA	-0.0027***	-0.0047***	0.0019***	-0.0051***	-0.0126***	0.0061***
Population	(0.0004)	(0.0005)	(0.0002)	(0.0008)	(0.0011)	(0.0006)

Panel B: Males

	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0004*	0.0000	-0.0004**	-0.0001	-0.0000	0.0000
Population	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0001)	(0.0001)
COFA	-0.0028***	-0.0040***	0.0011***	-0.0043***	-0.0072***	0.0024***
Population	(0.0004)	(0.0004)	(0.0002)	(0.0008)	(0.0008)	(0.0004)

*significant at the 10% level; **significant at the 5% level; ***significant at the 1% level

Notes: All estimations use 80,251 females and 76,693 males. All other notes per Table 2.

Table 6: Fixed Effects DD Estimates: Utilization by the Uninsured

	Under 65		Over 65	
	Inpatient	ER	Inpatient	ER
Japanese	0.0000*	0.0002***	-0.0000	0.0000
Population	(0.0000)	(0.0000)	(0.0000)	(0.0000)
COFA	0.0000	0.0016***	-0.0001	-0.0008*
Population	(0.0001)	(0.0002)	(0.0002)	(0.0004)

Notes: All outcomes are counts of admissions per patient/month that were charged to the individual. All other notes are per Tables 2 and 3.

Table 7: Fixed Effects DD Estimates: Effects on Infants

	Inpatient		ER		Inpatient Charges		ER Charges	
COFA	-0.0043***	-0.0033***	-0.0100***	-0.0106***	-99.4322***	-87.4223***	-21.7791***	-22.6053***
Population	(0.0003)	(0.0003)	(0.0007)	(0.0007)	(13.0656)	(11.9032)	(1.4574)	(1.4949)
COFA *		-0.0978***		0.0617***		-1147.4594**		86.9184***
Infant		(0.0151)		(0.0104)		(392.1023)		(15.9462)

Notes: All outcomes were reimbursed by Medicaid. The estimations were restricted to people under 65 years of age. All specifications correspond to those from Table 2 except that the odd columns include a dummy for being an infant and its interaction with the COFA/post variable. All other notes are per Table 2.

Table A1: Fixed Effects DD Estimates: Under 65 (Alternative Weights)

Panel A	Inpatient Admissions			ER Visits		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0002*	0.0001	-0.0003**	-0.0002	-0.0002*	0.0002
Population	(0.0001)	(0.0000)	(0.0001)	(0.0002)	(0.0001)	(0.0001)
COFA	-0.0023***	-0.0038***	0.0014***	-0.0041***	-0.0088***	0.0038***
Population	(0.0003)	(0.0003)	(0.0001)	(0.0005)	(0.0006)	(0.0003)

Panel B	Total Utilization			Total Charges		
Payer	Any	Medicaid	Private	Any	Medicaid	Private
Japanese	-0.0004*	-0.0001	-0.0001	-9.8978*	-1.8727	-4.6173
Population	(0.0002)	(0.0001)	(0.0001)	(4.2813)	(2.0709)	(3.2178)
COFA	-0.0065***	-0.0126***	0.0052***	-43.0653***	-106.3458***	50.8276***
Population	(0.0007)	(0.0008)	(0.0004)	(12.7479)	(12.2000)	(6.1824)

*significant at the 10% level; **significant at the 5% level; ***significant at the 1% level

Notes: Per Table 2. Additionally, the results in this table employ alternative weights based off of the 5-year ACS count of Micronesian in Hawaii plus its margin of error, which is 31,653.

Table A2: Fixed Effects DD Estimates: Under 65, Admissions and Charges in Logs

Payer	Any	Medicaid	Private	Any	Medicaid	Private
	Total Utilization (Logs)			Total Charges (Logs)		
Japanese	-0.00*	-0.00	-0.00	-0.00**	-0.00	-0.00
Population	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
COFA	-0.00***	-0.01***	0.00***	-0.05***	-0.10***	0.04***
Population	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)

*significant at the 10% level; **significant at the 5% level; ***significant at the 1% level

Notes: Per Table 2.

Table A3: Fixed Effects DD Estimates: Under 65 (with and without time dummies)

Payer	Any		Medicaid		Private	
			Inpatient Admissions			
Japanese	-0.0002*	-0.0002	0.0001	0.0001	-0.0003**	-0.0002**
Population	(0.0001)	(0.0001)	(0.0000)	(0.0000)	(0.0001)	(0.0001)
COFA	-0.0027***	-0.0028***	-0.0043***	-0.0044***	0.0015***	0.0015***
Population	(0.0003)	(0.0003)	(0.0003)	(0.0003)	(0.0001)	(0.0001)
			ER Visits			
Japanese	-0.0002	-0.0002	-0.0002*	-0.0002*	0.0002	0.0002
Population	(0.0002)	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
COFA	-0.0047***	-0.0047***	-0.0100***	-0.0100***	0.0043***	0.0043***
Population	(0.0006)	(0.0006)	(0.0007)	(0.0007)	(0.0003)	(0.0003)
Time Effects	No	Yes	No	Yes	No	Yes

*significant at the 10% level; **significant at the 5% level; ***significant at the 1% level

Notes: Per Table 2.

Figure 1: Utilization by Race

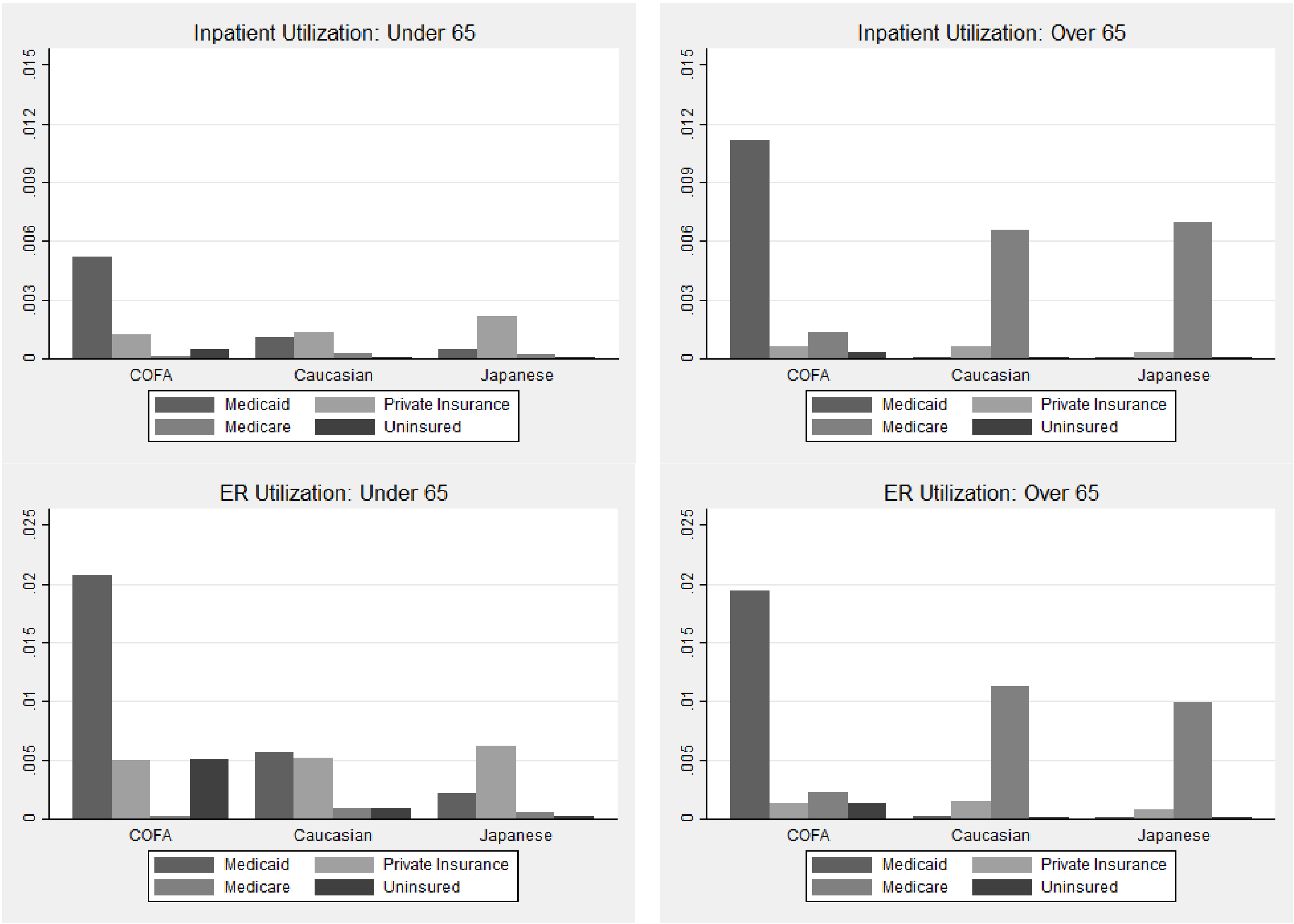

Figure 2: Event Analysis, Under 65

Figure 3: Effects by Age

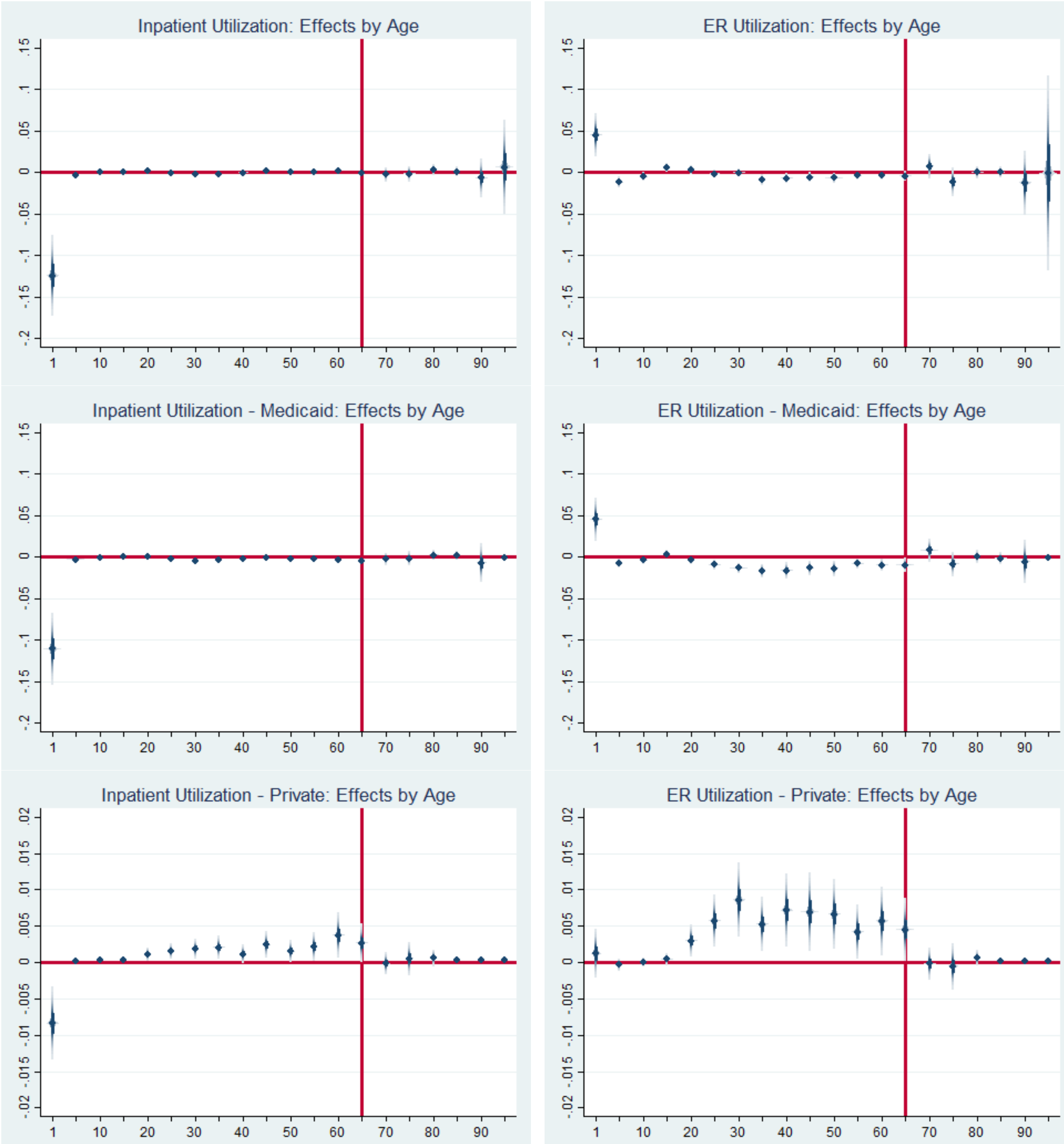

Figure 4a: Event Analysis, Under 65, Uninsured Utilization

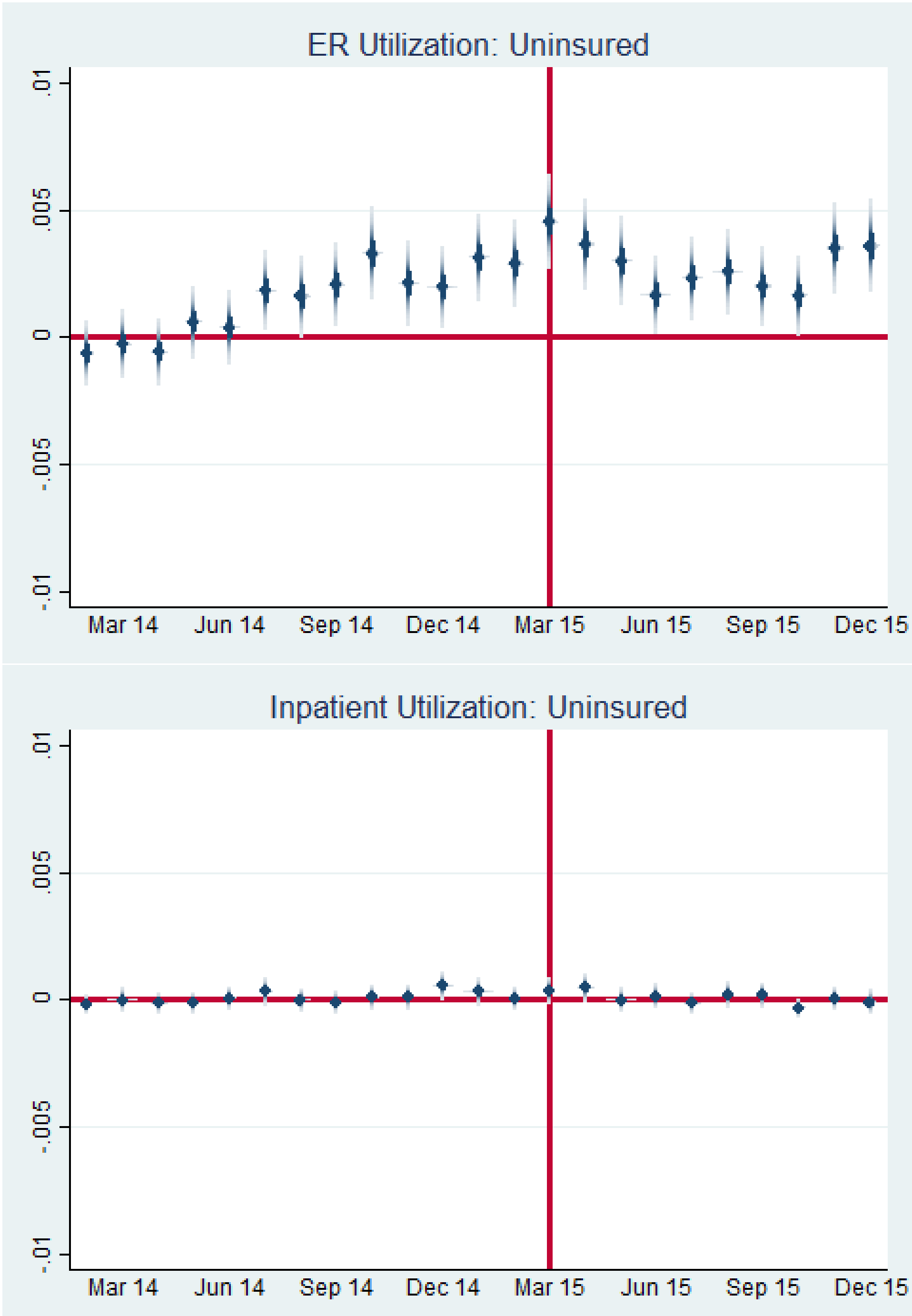

Figure 4b: Effects by Age, Uninsured Utilization

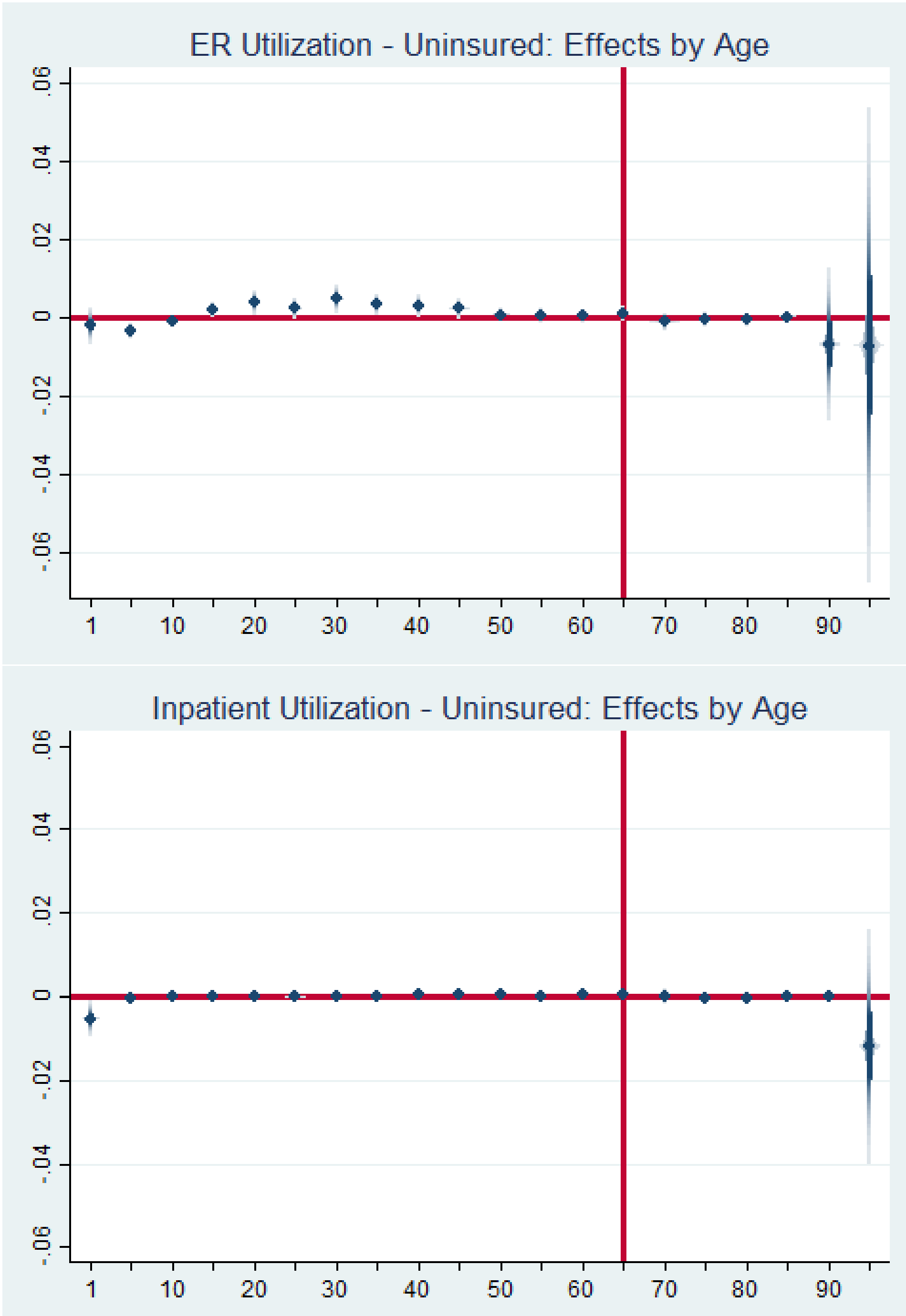

Figure A1: Effects by Age, Inpatient Utilization Funded by Medicaid
Excluding Infants