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

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October 24, 2018

Abstract

In March of 2015, the State of Hawaii stopped covering migrants from countries in the Compact of Free Association (COFA) in its Medicaid program forcing them to obtain private insurance in exchanges. Using administrative data, we show that Medicaid-funded hospitalizations and emergency room admissions declined in this population by 37 and 32%, respectively, after the expiration of Medicaid eligibility. Privately funded utilization did increase but not enough to offset declines in publicly-funded utilization resulting in a net decline. An exception to this is that infants born to COFA migrants were substantially more likely to be hospitalized after Medicaid benefits expired.

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

JEL Codes: I10, I14, J61

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We thank the University of Hawaii Economic Research Organization for providing funding and the Hawaii Health Information Corporation for providing the data. We would also like to thank seminar participants at the University of Hawaii Applied Micro Workshop for useful comments. All errors are our own.

I. Introduction

An important concern in health economics and public health is how insurance coverage affects the demand for health services. The reason for this is that those that lack financial resources are often the ones that are most in need of medical services. In the absence of adequate insurance, poorer populations might forgo necessary medical care. These concerns, especially in the United States, have been a driving force for the expansion of government-provided or government-subsidized health insurance. The Affordable Care Act (ACA) of 2010 established subsidies for the poorest 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 about 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).

Paradoxically, at a time when the United States was expanding insurance coverage for its poorest citizens, the State of Hawaii did the opposite for a small but vulnerable portion of its population. Up until March of 2015, the State of Hawaii enrolled migrants from countries in the Compact of Free Association (COFA) in its Medicaid program known as Medi-QUEST. COFA migrants are drawn primarily from the Republic of Palau, the Republic of the Marshall Islands, and the Federated States of Micronesia (FSM), three nation-states located in the middle and western Pacific Ocean.[1] At the time, the State of Hawaii had roughly 34,000 COFA migrants (commonly referred to as "Micronesians") most of whom were not US citizens but were guaranteed certain prerogatives such as access to public health insurance and free entry to the US due to the terms of the Compact. Access to federally-funded health insurance was suspended in 1996 due to welfare reform, but access to state-funded health insurance in various guises persisted in Hawaii due to the generosity of the State. However, this was suspended due to a

[1] The Compact of Free Association was signed for the Federated States of Micronesia and the Republic of the Marshall Islands in 1986 (in 1994 for Palau). Previously, these island groups were under part of the Trust Territory of the Pacific which was administered by the US 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 United States under a compact of free association. 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. Additionally, FSM citizens were allowed free entry into the United States at any time and were provided access to medical coverage such as Medicaid and other governmental services.

court ruling in April of 2014.[2] As a consequence, COFA migrants were denied access to Medicaid benefits in March of 2015 forcing them to purchase private insurance in the exchanges that were newly established as a consequence of the ACA, but they were not eligible for the Medicaid expansion (McElfish, et al. 2015). In this paper, we investigate how the expiration of these benefits impacted inpatient and emergency medical utilizations in this poorer population.

The expiration of public insurance benefits could have impacted the medical utilization of COFA migrants in two ways. The first is that it most likely increased the per-unit cost of services since Medicaid places heavy restrictions on co-payments and prohibits balance billing, the practice of providers charging patients for what insurers do not reimburse. Theoretically, this should 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). The second is that 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. This would also involve an increase in out-of-pocket expenses since many COFA migrants might have been uninsured after the expiration of their Medicaid benefits. This issue might have been especially important in this immigrant population where education levels and literacy rates are substantially lower than they are for natives. For example, Akee (2010) shows 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 10.3. At the same time, Baiker, et al. (2012) show that take-up rates of low-cost health insurance tend to be very low among the poor and less educated. This suggests that COFA migrants might have had a difficult time obtaining health insurance after the expiration of their Medicaid benefits.

[2] For details, we refer the reader to McElfish, et al. (2015).

In this paper, we employ administrative data on hospital discharges from Hawaii to estimate the effects of 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. Importantly, the data contain a unique patient identification number which enables us to track individual utilization over time. Based on 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 also contain an ethnicity variable. We employ data for three ethnicities: COFA migrants who are our treatment group, non-Hispanic whites who are our control group, and Japanese who are our placebo group. Omitted zeros for non-utilizers were dealt with by employing dummy observations and frequency weights that were constructed from population numbers from the American Community Survey (ACS).

To investigate the impact of the expiration of Medicaid benefits on utilization among COFA migrants, we employ 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 Medicaid after the conclusion of the State Medi-QUEST program for COFA migrants. In particular, Medicaid-funded ER and inpatient utilizations declined by 32 and 37%, respectively. This sharp reduction in utilization is consistent with other studies that have investigated the impact of the expiration of Medicaid benefits in Tennessee after it discontinued Medicaid benefits after a previous expansion in 2005 (e.g. DeLaire 2018, Tarazi 2017, Tello-Trillo 2016). At the same time, there was a large increase in the number of emergency and in-patient utilizations charged to private payers. This is important because it shows that there was a move towards private insurance among COFA migrants after their Medicaid benefits expired. However, the magnitude of this increase was smaller than the reduction in Medicaid-funded utilizations. As a result, on net, total inpatient and ER utilization declined which is consistent with some degree of moral hazard in this population.

Another important result in this study is that, while utilization in the COFA population declined overall, hospitalizations and emergency room admissions of infants born to COFA migrants *increased* as a consequence of the policy change. Due to data limitations, it is difficult for us to pin down the precise mechanism underlying this finding. However, it is consistent with a story in which pregnant mothers received less pre- and post-natal care once they left the state's Medicaid program which increased the need for medical care for their newborn children.

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. One perspective is that healthcare is completely determined by "needs" and so should not respond to price. However, since the RAND Health Insurance Experiment which we discussed previously, the economics profession has known this not to be the case. Indeed, the prevailing evidence suggests that the demand for medical care slopes downward so that there is moral hazard. The bulk of our results are consistent with this view.

However, a competing view is that the demand for medical care slopes *upwards* so that people that face the least cost sharing consume the least during the course of year (Einav and Finkelstein, 2018). The idea is that consuming health care that is ostensibly preventative in nature today can forestall the need for costlier care in the future.[3] Much of the discussion surrounding the ACA had this flavor with some policy makers speculating that expanded access to insurance would allow people to consume cheaper preventative care rather seek treatment in the ER.[4] In contrast to this commonly espoused view, evidence from the Oregon Health Insurance Experiment (Finkelstein et al, 2012) shows the opposite, namely, that randomized access to Medicaid *increases* ER utilization which is something that we also find. Importantly, however, we do find evidence of an increase in hospitalizations of infants born into COFA families *after* the expiration of Medicaid benefits. This is suggestive of an upward sloping demand curve for pregnant mothers and their children.

We conclude that these results provide additional evidence of the responsiveness of the demand for health services to the cost of those services. 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 does not compensate for the decline in utilizations that were previously financed by Medicaid. This suggests that there are now COFA migrants that are going without possibly important health care services as a result of the reduction in the program.

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

[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 of this in the case of Michigan's expansion of its Medicaid program after the implementation of the ACA.

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. Finally, we conclude.

II. Some Institutional Background

Publically-sponsored health care coverage for COFA migrants by the state of Hawaii has been subject to various state and federal policy decisions that have had very immediate effects on insurance coverage. These rapid changes have also led to general confusion in the COFA community about the actual health policies for which individuals are eligible. As argued above, this confusion can also itself be a barrier to insurance enrollment and to obtaining needed health care, particularly given the economic and social vulnerability of this community. Moreover, profound linguistic and cultural barriers to accessing government assistance are also a major impediment to obtaining insurance in this community. A brief overview of recent, relevant policies follows.

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. In this Act, immigrants to the US were made ineligible for federal Medicaid assistance without 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. Thus, most are not able to meet the five year waiting period to meet eligibility requirements. These COFA migrants are permanently classified as "non-qualified aliens." To make up for the shortfall in the wake of the 1996 welfare reform, starting in 1997, the state of Hawaii provided comprehensive health coverage for COFA residents under its state-funded Medicaid program (Rilkon et al., 2010). However, given that the agreement with COFA nations is a federal policy, the financial responsibility for providing these benefits has often been viewed as overly burdensome to the State in the absence of a significant increase in federal support (Hawaii DHS, 2009).

A seemingly important change to insurance coverage of COFA migrants happened in 2010 with the passage of the ACA. The ACA permitted COFA residents along with other lawfully present noncitizens were eligible to purchase health insurance through state health insurance exchange programs. However, Medicaid-ineligible noncitizens would not be eligible for federal

subsidies that would allow the individual to have no premiums. Instead, the ACA required Medicaid-ineligible noncitizens with incomes less than 100% 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, the State of Hawaii elected to cancel the Medicaid program eligibility for non-pregnant adult COFA residents due to budgetary shortfalls. Instead, the State created a limited medical assistance program called Basic Health Hawaii (BHH). Several court cases contesting this change in policy were filed. A lawsuit was filed and 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 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 who were not aged, blind or disabled (ABD) became ineligible for Medicaid benefits beginning in March of 2015 (Hawaii DHS, 2014; McElfish et al, 2015). However, legally residing children and pregnant women remained eligible for Medicaid from Children's Health Insurance Program Reauthorization Act (CHIPRA) and those who were ABD were able to receive the same level of benefits as those available under Medicaid. Beginning in March of 2015, Medicaid coverage effectively ended for COFA migrants in the State of Hawaii except for children, pregnant women, and people who were ADB.

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,

making this very different than Medicaid coverage.[5] However, Kaiser, which enrolled about half of the COFA migrants, waived these costs for those meeting eligibility requirements demonstrating financial need.

This policy change 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, the policy was confusing and enrollment on the exchange itself was confusing. Adding to this, the Hawaii Health Connector was troubled by technical challenges and in 2015, only a few months after the enrollment period, was closed down to be replaced by a federally-managed exchange. This meant that anyone who had been enrolled in the connector had to re-enroll in using Healthcare.gov. Once again, this caused confusion and necessitated outreach to the COFA community (Princeton, 2017). Unlike the Hawaii Health Connector, Healthcare.gov does not have linguistic outreach in COFA languages, likely adding even more challenges to this new enrollment (Princeton, 2017). [6]

III. Data

The data that we employ come from the Hawaii Health Information Corporation (HHIC), which is a private, not-for-profit organization based in Honolulu, Hawaii. HHIC collects data from hospitals in Hawaii and maintains these data in a large database. Its catchment area includes all of the hospitals in the State of Hawaii except for Tripler Army Medical Center. The data are nearly a census.

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 Japanese, Caucasian, and Micronesian people. In total, we used data on 409,556 specific utilizations. For our analysis, we only use utilizations for Hawaii residents *i.e.* people with

[5] It appears as if there were some practical difficulties with apply for the premium assistance program. For example, if the migrants do not properly file their 1095A tax forms, they are dis-enrolled. In addition, all correspondence is sent in English concerning the program is sent in English. Finally, each year a person applies to the premium assistant plan they must first get a denial letter from QUEST-Medicaid stating that they are ineligible for the QUEST program.

[6] In 2015 only, given this confusion, the State's Medicaid program did institute auto-enrollment so those who were being dropped and who had not chosen a plan were automatically paced into either a Kaiser or Hawaii Medical Service Association (HMSA) with an even split. While precise numbers are challenging to obtain, a recent policy analysis calculated that approximately 3,600 COFA Hawaii residents enrolled in coverage in Kaiser in 2015 and about 5,500 in HMSA (Princeton, 2017).

addresses in the State of Hawaii. These data include information on the type of discharge (*i.e.* inpatient or ER), admission and discharge dates, race (*e.g.* Micronesian, Japanese, or Caucasian), gender, age, payer type (*e.g.* Medicaid or private insurance), total billed, as well as the principal diagnosis and procedure codes. A critical feature of these data is that they include a synthetic master patient identification number which allows us to track 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 strongly indicates that no utilization took place in that month given the large catchment area of the HHIC data.

Descriptive statistics from the raw discharge data are reported in Table 1a. The bulk of our sample is Caucasian who constitute 65.6% of all utilizations in the discharge data. Next, we have that Japanese constitute 28.2% and that Micronesians comprise 6.3% of the discharge data. This sample has slightly more women (51.2%) than men (48.8%). Finally, most of the utilizations in our sample were for people on private plans (32.4%), followed by Medicare (28.3%), and then by Medicaid/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 charges exceeded one million dollars. This resulted in a final panel data set containing 205,688 individuals and 4,936,471 month/individual observations.

Using the HHIC data comes with an important sampling issue that requires some attention. By necessity, the HHIC data and the resulting panel that we just described only includes individuals of particular ethnicities that have had at least one admission to a hospital or an ER during 2014-2015. What this means is that we do not include people who had no such contact with the medical system during this time *i.e.* people who had no hospital or ER admissions 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.

To provide some idea of how the counts in the HHIC data correspond to the population counts in the State of Hawaii, we present population counts from the American Community Survey (ACS) in Table 1b. The counts from the ACS correspond to people who report Micronesian, Japanese, or White as one of their ethnicities. Accordingly, we account for the possibility that people can have multiple ethnicities which is a feature of the HHIC data. According to our estimates, the population counts of Micronesians, Japanese, and Whites are, respectively, 33976, 310595, and 604474. The corresponding counts in the HHIC data are 11530, 63174, and 131510 indicating that there are many missing zeros from our panel.

The solution to this is fairly simple. For each of the three ethnicities that we consider 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. Additional details can be found in the appendix.

Summary statistics on utilization and charges from the panel are reported in Tables 1c and 1d. All statistics use the frequency weights and so 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 all individuals in the top panel, people under 65 in the middle panel, and people over 65 in the bottom panel. Utilization and charges are broken down by inpatient and ER admissions. We also report statistics for all admissions which are under the heading, "all payers," those charged to Medicaid, those charged to private insurance, and those that were not charged to anyone which are under the heading, "uninsured."

In the top panel of Table 1c, we see that, on average, there were 0.0043 hospital and 0.0134 emergency room admissions per patient/month. This translates to a hospital 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 $181.17. The average amount charged to Medicaid was $33.95 and to private payers was $47.91. The remainder was paid by other payers such as Medicare.

In Table 1d, we report 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 main take-away from this table is that COFA migrants are much less healthy than the overall sample. For example, the mean of hospital admissions per patient/month among COFA migrants under 65 was 0.0061 whereas it was 0.0032 for the entire sample under age 65 in the previous table. Accordingly, the hospitalization rate for COFA migrants is double what is for the general population.

In Figure 1, we display histograms depicting total admissions per person/month charged to Medicaid, private insurance, Medicare, and the individual by ethnic group. The top panel shows inpatient admissions and the bottom panel shows ER admissions. The left panel shows utilizations for people under 65 and the right panel corresponds to people older than 65.

We see some salient patterns in the figure. 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 jump in charges to public insurance for people older than 65. 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 are admitted to the ER at much higher rates than the other two groups. Fourth, there is a much higher rate of uninsured COFA migrants than either the Japanese or Caucasians. This is particular clear in the bottom row of the figure where we look at ER admissions.

IV. Methods

To identify the effects of the move from Medicaid to private insurance 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 time t. A time period in our study is one of the 24 months between January 2014 and December 2015.

The main outcomes that we consider are the total number of hospital or ER admissions in a month and the total amount charged for these admissions. We further break these admissions down by those that were funded by Medicaid, those that were funded by private insurance, and those that were charged 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 in-patient or ER utilization in total as a result of the change in eligibility. We then separate out the outcome variable by whether the subsequent health services were paid by private insurance or Medicaid.

We employ three ethnic groups as treatment, control, and placebo groups. The treatment group is the COFA population which is identified as "Micronesian" in our data. The control group is Caucasian people. The placebo group is Japanese people.

For some outcome variable y_{it} , the main estimation equation that we employ can then be written 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 if the individual is Japanese, $COFA_i$ 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 time-invariant characteristics that might impact medical demand. Its inclusion eliminates the need to include any additional time-invariant characteristics (*e.g.* race dummies). The parameter π is the coefficient on the placebo and should be zero. Our parameter of interest is τ. This is 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 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.

Finally, we also estimate two 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. These estimations will also shed light on the parallel trends assumption. The second 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.

V. Results

Core Results

Our core results from the fixed effect DD model are reported in Table 2. The table reports the results of 18 estimations for people under 65 years of age. For each of the estimations, we report the coefficient on the post-dummy as well as the treatment effect on the COFA population and the placebo estimate corresponding to the interaction of the post and Japanese dummies.

In the first panel of Table 2, we look at the effects of the policy change on inpatient and ER admissions. 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 small negative impact on all inpatient admissions with a coefficient estimate of -0.0006 ($p<0.10$). Note that this effect is inclusive of those 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 admissions of 0.0019 ($p<0.001$) per person/month. These estimates imply that there were roughly 127 fewer hospital admissions and 400 fewer ER admissions among the COFA migrants over the 10 month period from March 1, 2015 to December 31, 2015.[7] Finally, because we see declines in both inpatient and ER admissions together, this is suggestive that the two types of utilizations are complements not substitutes as others have suggested.

Next, looking at admissions broken down by type of insurer, we see that the admissions funded by Medicaid declined whereas those funded by private insurance increased. However, the magnitudes of the former effects are larger than those for the latter effects which is what underlies the net negative impacts that we see in the first and fourth columns. In the second and fifth columns, we see that inpatient and ER admissions that were funded by Medicaid declined by 0.0019 ($p<0.01$) and 0.0062 ($p<0.01$) per patient/month. The means of inpatient and ER admissions among COFA migrants that were reimbursed by Medicaid in the pre-policy period were 0.0051 and 0.0192 in Table 1d. Accordingly, these effects amount to 37% and 32% decreases. In contrast, utilizations that were funded by private insurance and Medicaid both increased by 0.0013 ($p<0.01$) and 0.0035 ($p<0.01$). Using the means of 0.00055 and 0.00269 of Medicaid-funded inpatient and hospital admissions from Table 1d, we obtain that these are roughly 241% and 130% increases which is substantial. These large numbers are entirely attributable to the fact that the vast majority of COFA migrants were not enrolled in private insurance pre-policy. On the whole, this indicates that the policy worked as expected with a shift away from funding utilization of the COFA population with Medicaid towards funding it with private insurers. One important implication of this is that, since out-of-pocket costs for ER admissions tend to be higher for the private insurers in our data 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 and ER admissions. The first three columns of the panel, mechanically, are the sum of the impacts on inpatient and ER admissions from the first panel. In the first column, we see that, on net, utilization decreased by 0.0025 ($p<0.01$) admissions per person/month. The next

[7] To see this, note that there were roughly 21,187 COFA migrants under age 65 in Hawaii at this time. Accordingly, we will have that 0.0006*21,099*10=127.

two columns indicate that total utilization funded by Medicaid declined by 0.0081 ($p<0.01$) and those funded by private insurers increased by 0.0048 ($p<0.01$). The final three columns are analogous to the first three columns except that we use the log of total utilization as the dependent variable. We do this as a robustness check against potential outliers. All three estimates are significant at the 1% level. Note that because of the preponderance of zeros in the data, it is not appropriate to interpret these estimates as percentages.

The third panel reports the effects on total charges (of both inpatient and ER admissions) and the log of total charges. In the first column, we do not see any evidence that the level of total charges was impacted by the policy. However, in the fourth column, we see evidence that the policy impacted the log of total charges; the DD estimate is -0.02 ($p<0.01$). Breaking these charges down by those funded by Medicaid and private insurance, we see that charges to Medicaid declined by $60.87 ($p<0.01$) per person/month in the second column and those charged to private insurers, in the third column, increased by $52.35 ($p<0.01$). Using the figures from Table 1d, these correspond a 35% decline and a 253% increase, respectively. Finally, the estimates using log charges in the final three columns of this panel are all significant suggesting that our results are not driven by outliers.

The placebo interactions in Table 2 are by-and-large not significant. Of the 18 separate estimations, only two of the placebo interactions are significant but only at the 10% level. 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 Table 3, we estimate the same models as in the first panel of Table 2 using inpatient and ER admissions 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 utilizations funded by private payers increase by 0.0007 ($p<0.10$) per person/month and ER admissions also funded by private payers increase by 0.0023 ($p<0.001$) per person/month. Similarly, we see that inpatient admissions funded by Medicaid decline by 0.0011 ($p<0.10$) and ER admissions funded by Medicaid decline by 0.0034 ($p<0.05$) per person/month.

In the second panel 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 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. On the whole, we see few impacts of the policy change in this older population. First, we see no effects for any admissions that were charged to Medicaid. We do, however, see modest effects on admissions that were charged to private payers, although relative to the corresponding effects in Table 2, these estimates are attenuated and tend to have higher p-values associated with the estimates. For example, in the third column of the first panel, we see that the policy change increased hospital admissions charged to private payers by 0.0005 ($p<0.10$) per person/month. By way of comparison, the analogous estimate from Table 2 is 0.0013 ($p<0.01$). Similarly, in the population older than 65, we see that the policy change increased ER admissions charged to private payers by 0.0006 ($p<0.10$) whereas the effect in the population under 65 was 0.0035 ($p<0.01$). The null effects for Medicaid-reimbursed admissions suggest that older COFA migrants continued to receive care from public insurance both before and after the policy change. However, the modest increase in admissions charged to private payers for the older population does indicate that the policy also impacted older COFA migrants.

This could have happened for two reasons. First, younger COFA migrants who enrolled in private insurance may have claimed older relatives or spouses as dependents on their policies. Second, the push to enroll COFA migrants in private plans after the policy's implementation may have also caused older migrants to enroll.

Event Analysis

In Figure 2, we report the results of an event analysis. The figure contains six graphs which display coefficients on the interactions between 24 month dummies and the COFA migrant dummy from a model that is otherwise similar to equation (1). The first row displays figures for inpatient and ER admissions funded by any payer. The second row displays figures for admissions funded by Medicaid and the final row displays admissions funded by private insurance. All six figures include a horizontal line at zero and a vertical line corresponding to 1 March 2015.

We see the following results. First, the interactions, on the whole, are not significantly different from zero in all six graphs prior to 1 March 2015. This provides some 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 admissions charged to Medicaid does show a dip in the pre-period. This result may indicate that the impending change in state coverage of Medicaid for COFA immigrants produced a response in emergency room visits prior to the actual change in the policy. 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 admissions charged to any payer does show a significant decrease in the post-period relative to the pre-period suggesting that there was a net decline in ER admissions attributable to the policy change. Fourth, in the second row of Figure 1, we see that both inpatient and ER admissions charged to Medicare declined after the policy's implementation. Finally, the third row indicates utilizations that were charged to private insurers displayed no trends in the pre-period but did increase substantially in the post-period.

Results by Age and Gender

In Figure 3, we investigate how the effects of the policy varied by age. We estimated a variation of equation (1) in which we included a set of age dummies and their interactions with the COFA dummy. 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 65.

The first row displays results for inpatient and ER admissions 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 positive effect. These age effects are very tightly estimated until 65 after which the confidence bands around them become substantially larger. The large positive effect that we see on infants underscores that the results from Table 2 mask some important heterogeneity. In particularly, while it does appear that the policy reduced most utilization, this was not the case for the youngest Micronesians. The second figure in the row shows a net decrease in ER admissions 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 admissions occurs for people between the ages of 30 and 65. Once again, we see an increase in ER admissions for Micronesian infants after the policy's implementation. On the whole, this suggests that the health of the very young may have been adversely impacted by the policy change.

In the second row of Figure 3, we display results for inpatient and ER admissions charged to Medicaid. The two graphs show that, overall, the policy reduced these ER admissions until age 65 whereupon the effects are not significantly different from zero. The effects for ER admissions are particularly pronounced in the second column. Importantly, the two figures in this row show that ER and inpatient admissions charged to Medicaid increased substantially for infants which is what underlies the corresponding increases that we saw in the first row of the figure. It is important to emphasize that the positive effects on infants are very large relative to the negative effects on other COFA migrants under age 65. However, these negative effects are in fact significantly different from zero. 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. This figure clearly shows that there were negative and significant effects on Medicaid-funded hospital admissions.

Prima facie, the increase in hospitalizations and ER admissions for infants is a puzzle since the policy moved COFA migrants off of Medicaid towards private insurance. One possible explanation for this finding is that pregnant mothers received worse pre- and post-natal care after the policy. However, one puzzle (per the discussion earlier) is that the pregnant women were still covered by Medicaid after March, 2015. Nevertheless, this aspect of the new policy may

have not been communicated well. As a potential result, this could have resulted in worse health outcomes for their children. If a health issue subsequently arose for these children, who may have been uninsured, the hospital may have billed it to Medicaid.

The final row of Figure 2 displays the results for admissions charged to private payers. Both figures in this row show modest increases in both inpatient and ER utilizations up to age 65 upon which the effects become null. Importantly, the magnitude of the effects for ER admissions 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.

In Table 5, we look at the impacts across genders. We report DD estimates of the effects of the policy on inpatient and ER admissions 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 females were impacted more than males. For example, female COFA migrants saw a decline in inpatient admissions charged to Medicaid of 0.0023 ($p<0.01$) per person/month, whereas males saw a decline of 0.0013 ($p<0.01$) per person/month. The corresponding estimates for ER admissions charged to Medicaid are -0.0088 ($p<0.01$) for females and -0.0035 ($p<0.01$) for males which again indicates an impact on females that is about twice as large as it is for males.

Effects on the Uninsured

We conclude by looking at how the uninsured were impacted by the policy change. Ostensibly, the State of Hawaii intended to smoothly transition COFA migrants from Medicaid to subsidized private insurance. However, it was widely suspected that many migrants did not enroll in private insurance in a timely manner. To shed light on this, we estimate our fixed effects DD model using inpatient and ER admissions 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 admissions for people who are both under and over 65. We see impacts for COFA migrants under 65. Uninsured inpatient admissions increased by 0.0001 per

person/month and uninsured ER admissions increased by 0.0014 ($p<0.01$) per person/month. Note that the impact on uninsured ER admissions is roughly half of the impact on ER admissions charged to private insurers of 0.0035 from Table 2. In addition, the impact on ER admissions charged to Medicaid in the same table was -0.0062. Accordingly, COFA migrants insured roughly 25% of the decline in ER admissions induced by the policy with uninsured visits to ER. The last two columns show no impacts on uninsured admissions among older COFA migrants.

In Figures 4a and 4b, we report 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 admissions. However, this increase pre-dates 1 March 2015 which indicates that there may have been some anticipation of the policy change. However, it is important to note that the plot of Medicaid-funded ER admissions from Figure 2 also shows a decline prior to the policy's implementation. This suggests that prior to the implementation of the policy there was a substitution away from Medicaid-funded ER admissions towards uninsured admissions even when this was not necessary. This indicates a possibly haphazard implementation of the policy. Second, in Figure 4b, we see that the bulk of the impacts on uninsured ER admissions occurred for adults younger than 65. Note that we do not see large impacts on uninsured ER admissions for infants. There were no impacts on COFA migrants older than 65. Finally, the effects on uninsured hospital admissions were relatively smaller.

Effects on the Infants

Finally, we further explore how the policy change impacted the youngest Micronesians. 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 and ER admissions, and inpatient 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.

The results indicate that the effects on the policy on infants were enormous. Looking at the first two columns where the dependent variable was inpatient admissions, we see that the raw impact of the policy was -0.0019 ($p<0.01$) in the first column. In the second column, the direct impact increases to -0.0029 ($p<0.01$), but the interaction with the infant dummy is 0.0605 ($p<0.01$). Accordingly, the net effect on infants was 0.0605-0.0029=0.058 inpatient admissions per person/month. The ratio of the magnitudes of direct effect and its interaction in the second column is 21. We see a similar phenomenon in the third and fourth columns where ER admissions are the dependent variable. The fourth column shows a direct effect of the policy on ER admissions of -0.0075 ($p<0.01$), but the interaction with the infant dummy is 0.0736 ($p<0.01$). This ratio is about 10. The net effect on ER admissions for infants is 0.0661 ER admissions.

The final four columns of the table report the effects on charges. These results indicate that the policy increased charges for inpatient admissions of Micronesian children by 1554.37-71.41= 1482.96 dollars per person/month. The corresponding number for ER admissions was 108.88-17.01=91.87 dollars per person/month.

VI. Conclusions

In this paper, we investigated the effects of removing a large swath of the Micronesian population in the State of Hawaii from Medicaid on their medical utilization. To do this, we employed a large administrative database that essentially constitutes a complete census of all inpatient and emergency room utilizations during 2014 and 2015. Difference-in-difference models indicate that the expiration of Medicaid benefits decreased Medicaid-funded inpatient and emergency room utilizations by 37% and 32%, respectively. Privately-funded utilizations increased by 241% (inpatient) and 130% (emergency room). On net, the magnitudes of the publicly-funded utilizations did not make up for the decline in Medicaid-funded utilizations resulting in a net decline in utilization after the expiration of Medicaid benefits. Importantly, some of the shortfall in Medicaid-funded utilizations was made up for by utilizations of uninsured patients. Finally, despite our finding that utilization on net decreased, we did show that Micronesian infants were more likely to be hospitalized after Medicaid benefits expired. This was the case despite the fact that the State of Hawaii made special arrangements to continue Medicaid benefits for pregnant Micronesian women.

References

Akee, R. (2010). Who leaves? Deciphering immigrant self-selection from a developing country. *Economic Development and Cultural Change*, *58*(2), 323-344.

Aviva Aron-Dine, Liran Einav, and Amy Finkelstein. 2013. "The RAND Health Insurance Experiment, Three Decades Later." *Journal of Economic Perspectives* Volume 27, Number 1 Winter: 197–222.

Baicker, K., Congdon, W. J., & Mullainathan, S. (2012). Health insurance coverage and take-up: Lessons from behavioral economics. *The Milbank Quarterly*, *90*(1), 107-134.

Card, D., Dobkin, C., & Maestas, N. (2008). The impact of nearly universal insurance coverage on health care utilization: evidence from Medicare. *American Economic Review*, *98*(5), 2242-58.

Cohen RA, Zammitti EP, Martinez ME. Health insurance coverage: Early release of estimates from the National Health Interview Survey, 2016. National Center for Health Statistics. May 2017. Available from: https://www.cdc.gov/nchs/nhis/ releases.htm.

Cutler, D. M., & Zeckhauser, R. J. (2000). The anatomy of health insurance. In *Handbook of health economics* (Vol. 1, pp. 563-643). Elsevier.

DeLeire, T. (2018). *The Effect of Disenrollment from Medicaid on Employment, Insurance Coverage, Health and Health Care Utilization* (No. w24899). National Bureau of Economic Research.

Einav, L., & Finkelstein, A. (2018). Moral Hazard in Health Insurance: What We Know and How We Know It. *Journal of the European Economic Association*, *16*(4), 957-982.

Finkelstein, Amy, Sarah Taubman, Bill Wright, Mira Bernstein, Jonathan Gruber, Joseph P. Newhouse, Heidi Allen, Katherine Baicker, and the Oregon Health Study Group. 2012. "The Oregon Health Insurance Experiment: Evidence from the First Year." *Quarterly Journal of Economics* 127(3): 1057–1106.

Goldman, D., & Philipson, T. J. (2007). Integrated insurance design in the presence of multiple medical technologies. *American Economic Review*, *97*(2), 427-432.

Hawaii Department of Human Services (DHS). (2009). DHS Moving Non-Citizen Adults into new Medical Assistance Program, 'Basic Health Hawai'i' on September 1. Hawaii Department of Human Services. News Release. July 28, 2009.

Hawaii Department of Human Services. COFA Residents and Health Care Assistance in Hawaii. November 2014. Available from: https://humanservices.hawaii.gov/wp-content/uploads/2014/11/COFA-Background-Memo.pdf

Manning, Willard G., Joseph P. Newhouse, Naihua Duan, Emmett B. Keeler, Arleen Leibowitz, and M. Susan Marquis. 1987. "Health Insurance and the Demand for Medical Care: Evidence

from a Randomized Experiment." *American Economic Review* 77(3): 251–77.

McElfish, P.A., Hallgren, E. and Yamada, S., 2015. Effect of US health policies on health care access for Marshallese migrants. *American journal of public health*, *105*(4), pp.637-643.

Joseph P. Newhouse, Rand Corporation. Insurance Experiment Group, & Insurance Experiment Group Staff. (1993). *Free for all?: lessons from the RAND health insurance experiment*. Harvard University Press.

Pauly, M. V. (1968). The economics of moral hazard: comment. *The American Economic Review*, 531-537.

Princeton University. Hawaii's COFA Islanders: Improving Health Access and Outcomes. A report for the Governor of Hawaii, David Ige. January 2017. Available from wws.princeton.edu/sites/.../WWS%20591e%20ACA%20Final%20Report%202017.pdf

Rilkon S, Alik W, Hixon A, Palafox N. (2010). The "Compact Impact" in Hawai'i: Focus on health care. *Hawaii Medical Journal*, 69 (6); 7-12.

Schoen, C., Doty, M.M., Robertson, R.H. and Collins, S.R., 2011. Affordable Care Act reforms could reduce the number of underinsured US adults by 70 percent. *Health Affairs*, *30*(9), pp.1762-1771.

Shigeoka, Hitoshi. 2014. The Effect of Patient Cost Sharing on Utilization, Health, and Risk Protection. *American Economic Review* 104:7, 2152-2184.

Tarazi, W. W., Green, T. L., & Sabik, L. M. (2017). Medicaid disenrollment and disparities in access to care: evidence from Tennessee. *Health services research*, *52*(3), 1156-1167.

Tello-Trillo, D. S. (2016, June). Effects of Losing Public Health Insurance on Healthcare Access, Utilization and Health Outcomes: Evidence from the TennCare Disenrollment. In *6th Biennial Conference of the American Society of Health Economists, Ashecon*.

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

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

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 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.

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%]

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

	HHIC	ACS	HHIC/ACS
Micronesian	11530	33976	34%
Japanese	63174	310595	22%
White	131510	604474	29%

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

Table 1c: Descriptive Statistics from Panel Data

	Any Payer	Medicaid	Private	Uninsured
		All		
Inpt. Admissions	0.00431	0.00086	0.00134	0.00006
	(0.0691)	(0.0311)	(0.0380)	(0.0082)
ER Admissions	0.01355	0.00412	0.00446	0.00069
	(0.1391)	(0.0825)	(0.0731)	(0.0291)
Total Charges	181.17	33.95	47.91	2.89
	(4101.91)	(1819.59)	(2028.17)	(309.47)
Inpt. Charges	149.27	25.68	37.52	1.40
	(4058.61)	(1794.52)	(2007.27)	(293.90)
ER Charges	31.89	8.27	10.39	1.49
	(413.14)	(213.27)	(217.28)	(88.32)
		Under 65		
Inpt. Admissions	0.00320	0.00108	0.00159	0.00008
	(0.0594)	(0.0348)	(0.0414)	(0.0090)
ER Admissions	0.01378	0.00532	0.00547	0.00088
	(0.1409)	(0.0936)	(0.0810)	(0.0328)
Total Charges	121.82	41.77	54.21	3.45
	(3327.11)	(2000.25)	(2120.59)	(325.78)
Inpt. Charges	91.92	31.19	41.67	1.58
	(3283.15)	(1971.56)	(2097.47)	(307.49)
ER Charges	29.90	10.58	12.54	1.87
	(393.68)	(240.63)	(237.26)	(98.37)
		Over 65		
Inpt. Admissions	0.00781	0.00016	0.00054	0.00002
	(0.0932)	(0.0135)	(0.0244)	(0.0048)
ER Admissions	0.01283	0.00035	0.00127	0.00010
	(0.1333)	(0.0259)	(0.0390)	(0.0111)
Total Charges	367.82	9.36	28.10	1.12
	(5906.03)	(1067.69)	(1705.05)	(251.37)
Inpt. Charges	329.64	8.37	24.50	0.82
	(5859.47)	(1059.79)	(1692.49)	(246.34)
ER Charges	38.18	0.99	3.60	0.30
	(469.06)	(79.76)	(136.32)	(43.56)

Notes: Reports means and standard deviations in parentheses.

Table 1d: Descriptive Statistics from Panel Data: COFA Migrants

	Any Payer	Medicaid	Private	Uninsured
		All		
Inpt. Admissions	0.00638	0.00532	0.00053	0.00036
	(0.08253)	(0.0754)	(0.0239)	(0.0191)
ER Admissions	0.02589	0.01898	0.00259	0.00350
	(0.1750)	(0.1511)	(0.0546)	(0.0633)
Total Charges	241.65	194.94	20.53	15.75
	(4969.58)	(4654.99)	(1206.30)	(666.23)
Inpt. Charges	191.02	159.49	14.68	8.21
	(4930.68)	(4625.16)	(1189.60)	(632.56)
ER Charges	50.63	35.45	5.85	7.55
	(462.03)	(371.62)	(165.41)	(192.01)
		Under 65		
Inpt. Admissions	0.00608	0.00505	0.00055	0.00036
	(0.0803)	(0.0733)	(0.0242)	(0.01931)
ER Admissions	0.02617	0.01918	0.00269	0.00364
	(0.1762)	(0.1519)	(0.0555)	(0.0646)
Total Charges	219.45	174.21	20.66	15.88
	(4735.47)	(4409.44)	(1219.14)	(657.27)
Inpt. Charges	169.59	139.69	14.71	8.10
	(4697.40)	(4380.66)	(1203.84)	(621.98)
ER Charges	49.86	34.52	5.94	7.78
	(451.24)	(359.79)	(160.61)	(194.60)
		Over 65		
Inpt. Admissions	0.01129	0.00977	0.00033	0.00026
	(0.1121)	(0.1046)	(0.0183)	(0.0161)
ER Admissions	0.01942	0.01571	0.00093	0.00119
	(0.1519)	(0.1366)	(0.0350)	(0.0355)
Total Charges	610.94	539.81	18.37	13.56
	(7896.86)	(7653.90)	(967.92)	(800.66)
Inpt. Charges	547.66	489.00	14.07	9.89
	(7845.60)	(7610.44)	(920.74)	(787.99)
ER Charges	63.29	50.81	4.30	3.67
	(614.27)	(530.90)	(231.01)	(142.11)

Notes: Reports means and standard deviations in parentheses. All descriptive statistics correspond to the period prior to March 1, 2015.

Table 2: Fixed Effects DD Estimates: Under 65

Payer	Any	Medicaid	Private	Any	Medicaid	Private
		Inpatient			ER	
JP	-0.0001	-0.0000	0.0000	-0.0002	-0.0002	0.0002*
	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0001)	(0.0001)
COFA	-0.0006*	-0.0019***	0.0013***	-0.0019***	-0.0062***	0.0035***
	(0.0002)	(0.0002)	(0.0002)	(0.0005)	(0.0006)	(0.0003)
		Total Utilization			Log Total Utilization	
JP	-0.0004	-0.0003	0.0003	-0.0003	-0.0001	0.0001
	(0.0002)	(0.0001)	(0.0002)	(0.0001)	(0.0001)	(0.0001)
COFA	-0.0025***	-0.0081***	0.0048***	-0.0016***	-0.0052***	0.0031***
	(0.0006)	(0.0008)	(0.0005)	(0.0003)	(0.0005)	(0.0003)
		Total Charges			Log Total Charges	
JP	-3.36	-2.47	1.93	-0.00*	-0.00	0.00
	(5.13)	(2.20)	(3.49)	(0.00)	(0.00)	(0.00)
COFA	1.45	-60.87***	52.35***	-0.02***	-0.06***	0.03***
	(13.31)	(11.89)	(7.63)	(0.00)	(0.01)	(0.00)

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

Notes: All estimations use 154,395 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. JP is the interaction of the post dummy with the Japanese dummy. COFA is the interaction of the post dummy with the COFA dummy. Standard errors adjust for clustering on individuals.

Table 3: Alternative Estimations: Under 65

Payer	Any	Medicaid	Private	Any	Medicaid	Private
		Inpatient			ER	
			First Differenced-Fixed Effects			
JP	0.0000	0.0000	0.0001	-0.0004	-0.0003	-0.0001
	(0.0002)	(0.0001)	(0.0002)	(0.0005)	(0.0003)	(0.0003)
COFA	-0.0003	-0.0011*	0.0007*	0.0000	-0.0034**	0.0023***
	(0.0007)	(0.0006)	(0.0003)	(0.0015)	(0.0012)	(0.0006)
			OLS			
JP	-0.0001	-0.0000	0.0000	-0.0002	-0.0002	0.0003
	(0.0002)	(0.0001)	(0.0001)	(0.0004)	(0.0002)	(0.0002)
COFA	-0.0006*	-0.0019***	0.0012***	-0.0016**	-0.0060***	0.0035***
	(0.0003)	(0.0002)	(0.0001)	(0.0006)	(0.0006)	(0.0004)
			Fixed Effects			
JP	-0.0001	-0.0000	0.0000	-0.0002	-0.0002	0.0002*
	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0001)	(0.0001)
COFA	-0.0006*	-0.0019***	0.0013***	-0.0019***	-0.0062***	0.0035***
	(0.0002)	(0.0002)	(0.0002)	(0.0005)	(0.0006)	(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

Payer	Any	Medicaid	Private	Any	Medicaid	Private
		Inpatient			ER	
JP	0.0000	0.0000	0.0001*	0.0000	0.0001*	0.0003**
	(0.0002)	(0.0000)	(0.0001)	(0.0003)	(0.0001)	(0.0001)
COFA	-0.0006	-0.0013	0.0005*	-0.0007	-0.0006	0.0006*
	(0.0011)	(0.0010)	(0.0002)	(0.0017)	(0.0015)	(0.0003)
		Total Utilization			Log Total Utilization	
JP	0.0001	0.0001	0.0004**	0.0001	0.0001**	0.0003**
	(0.0005)	(0.0001)	(0.0002)	(0.0003)	(0.0000)	(0.0001)
COFA	-0.0013	-0.0018	0.0011*	-0.0010	-0.0014	0.0008**
	(0.0022)	(0.0021)	(0.0004)	(0.0014)	(0.0013)	(0.0003)
		Total Charges			Log Total Charges	
JP	3.35	-0.12	2.08	0.00	0.00**	0.00**
	(12.63)	(1.53)	(3.44)	(0.00)	(0.00)	(0.00)
COFA	-17.55	-33.77	13.21	-0.01	-0.02	0.01**
	(78.92)	(76.15)	(9.99)	(0.02)	(0.01)	(0.00)

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

Notes: All estimations use 54,331 individuals (including the dummy observations). All other notes are per Table 2.

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

Payer	Any	Medicaid	Private	Any	Medicaid	Private
		Inpatient			ER	
			Females			
JP	-0.0001	0.0000	0.0001	-0.0005	-0.0005*	0.0004*
	(0.0002)	(0.0001)	(0.0001)	(0.0003)	(0.0002)	(0.0002)
COFA	-0.0008*	-0.0023***	0.0016***	-0.0027***	-0.0088***	0.0049***
	(0.0003)	(0.0004)	(0.0002)	(0.0007)	(0.0011)	(0.0006)
			Males			
JP	-0.0001	-0.0001	0.0000	0.0000	0.0000	0.0001
	(0.0002)	(0.0001)	(0.0001)	(0.0003)	(0.0002)	(0.0001)
COFA	-0.0003	-0.0013***	0.0010***	-0.0011	-0.0035***	0.0019***
	(0.0003)	(0.0003)	(0.0002)	(0.0007)	(0.0007)	(0.0004)

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

Notes: All estimations use 79,039 females and 75,356 males. All other notes are per Table 2.

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

	Under 65		Over 65	
	Inpatient	ER	Inpatient	ER
JP	0.0000*	0.0002***	0.0000	0.0000*
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
COFA	0.0001	0.0014***	0.0000	-0.0003
	(0.0000)	(0.0002)	(0.0001)	(0.0003)

Notes: All outcomes are counts of admissions per person/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	-0.0019***	-0.0029***	-0.0062***	-0.0075***	-45.68***	-71.41***	-15.19***	-17.01***
	(0.0002)	(0.0003)	(0.0006)	(0.0007)	(11.38)	(11.60)	(1.40)	(1.54)
COFA *		0.0605***		0.0736***		1554.37***		108.88***
Infant		(0.0035)		(0.0057)		(255.66)		(9.09)

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.

Figure 1: Utilization by Race

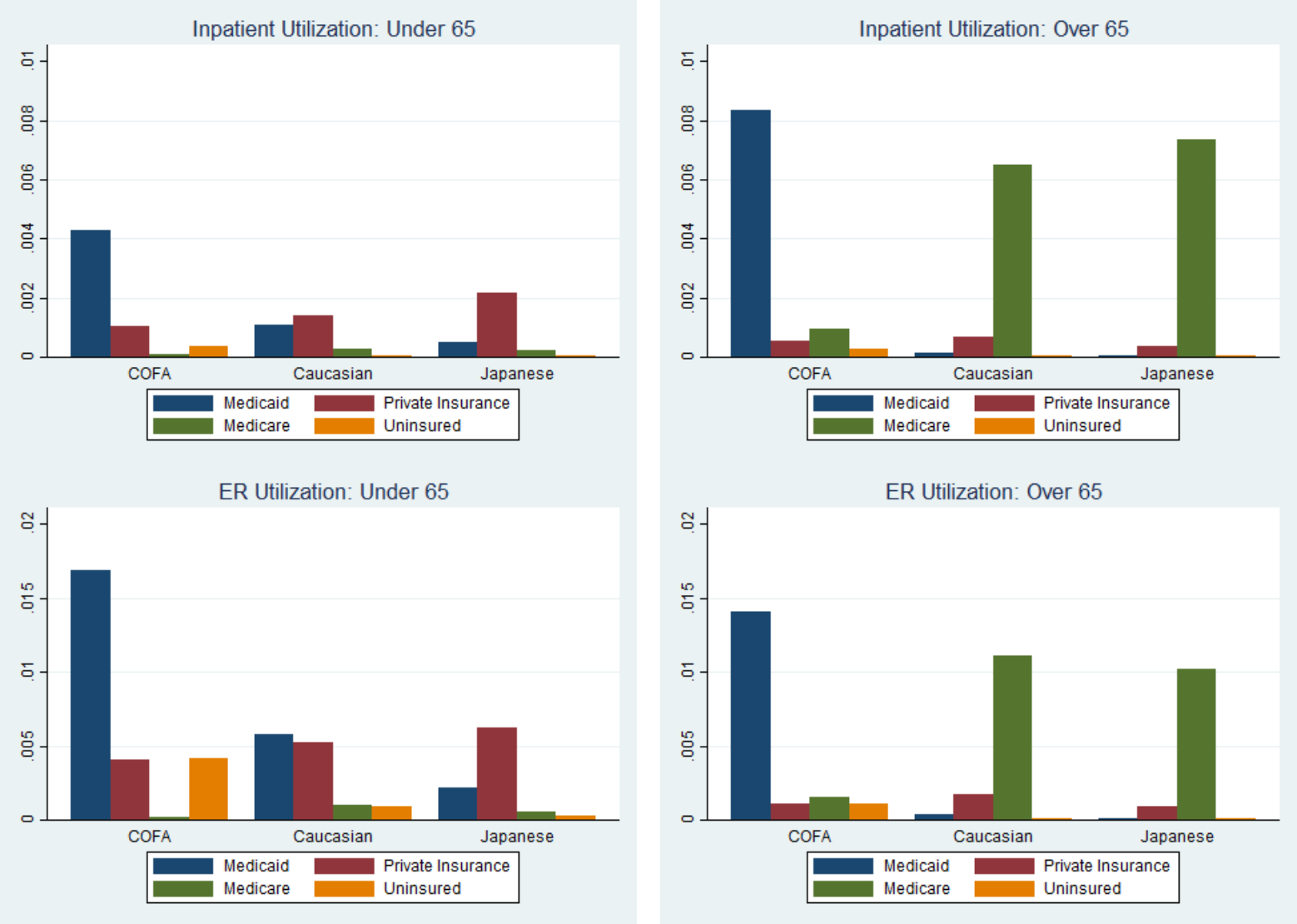


Figure 2: Event Analysis, Under 65

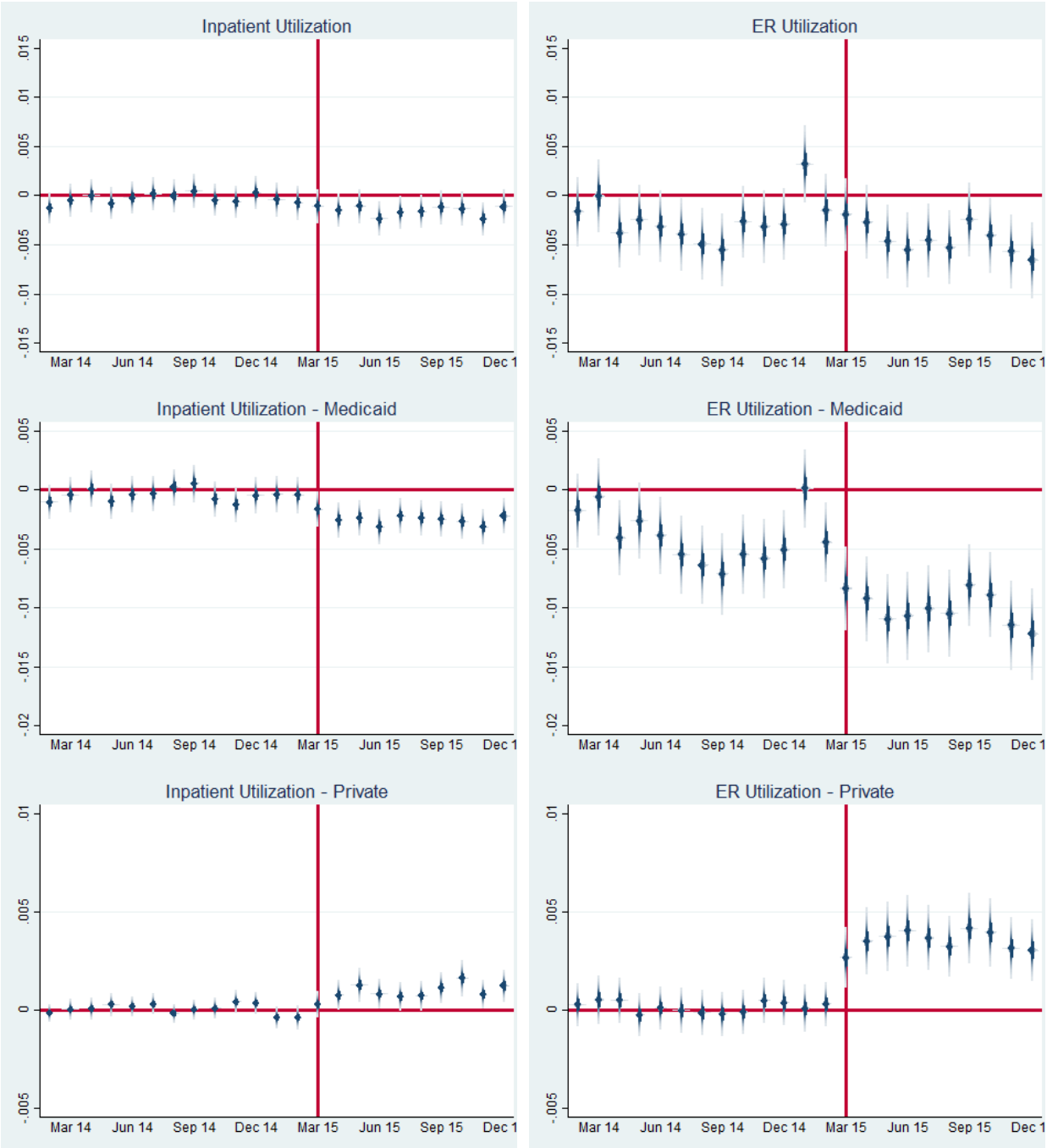


Figure 3: Effects by Age

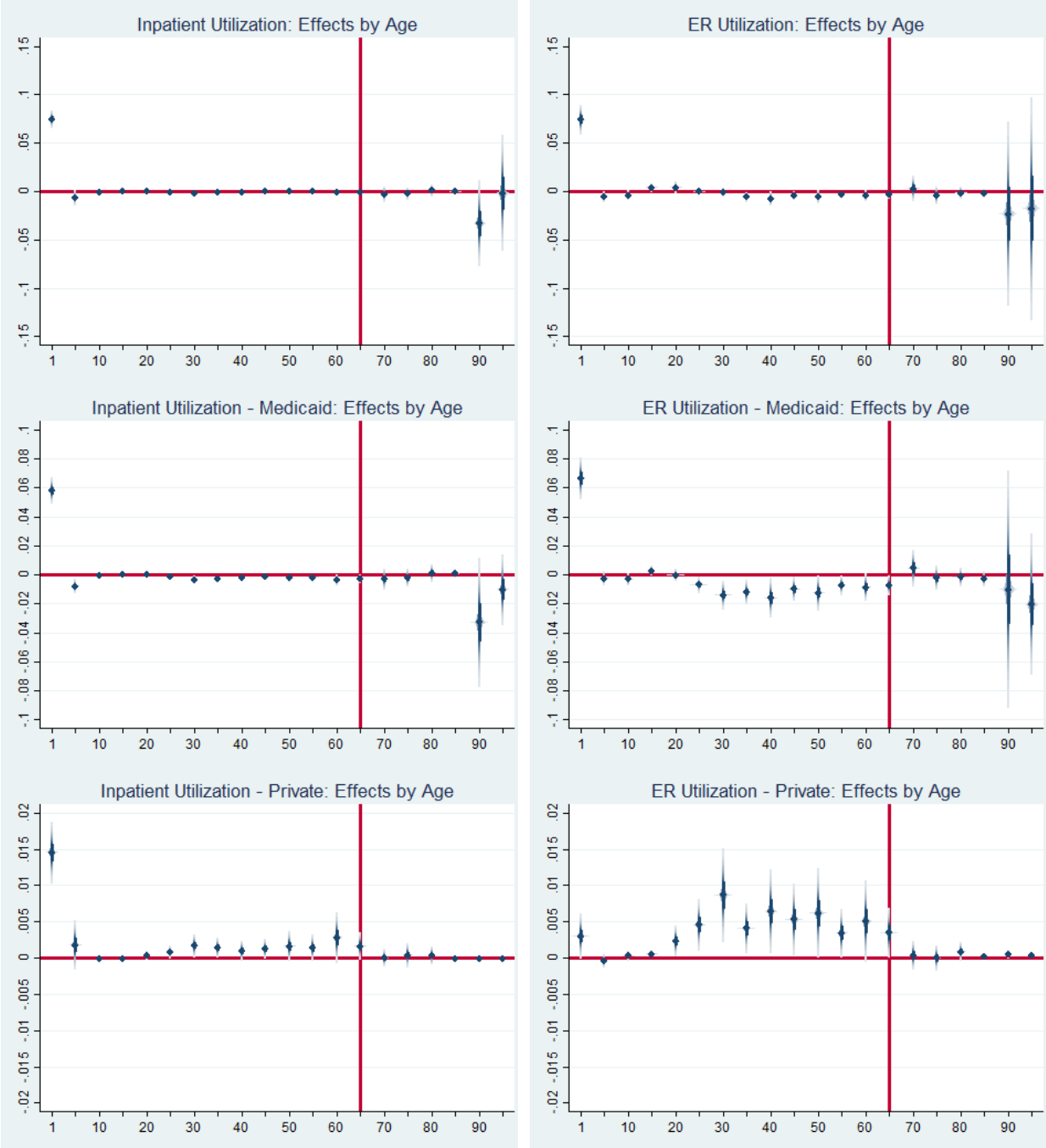


Figure 4a: Event Analysis, Under 65, Uninsured Utilization

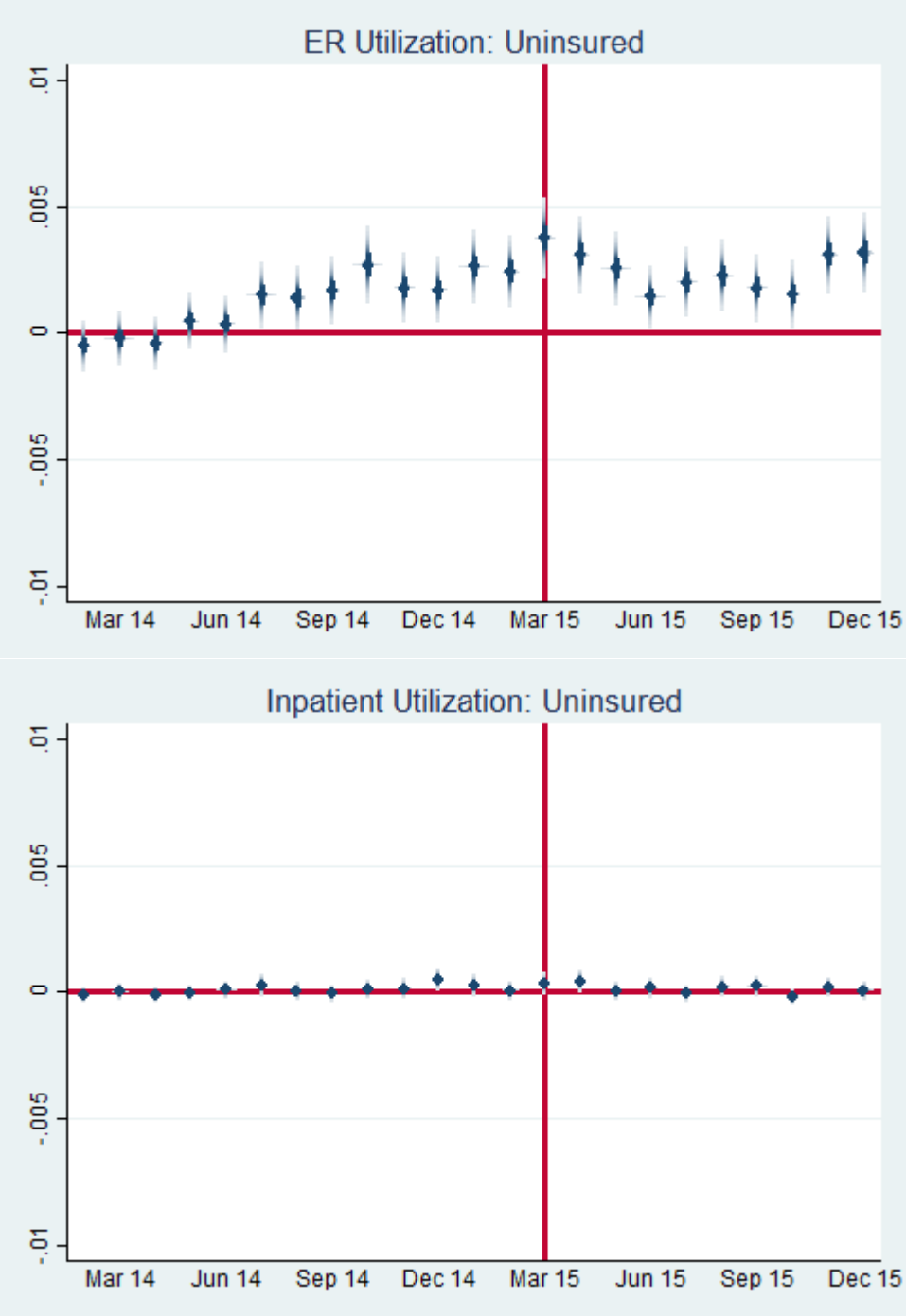


Figure 4b: Effects by Age, Uninsured Utilization

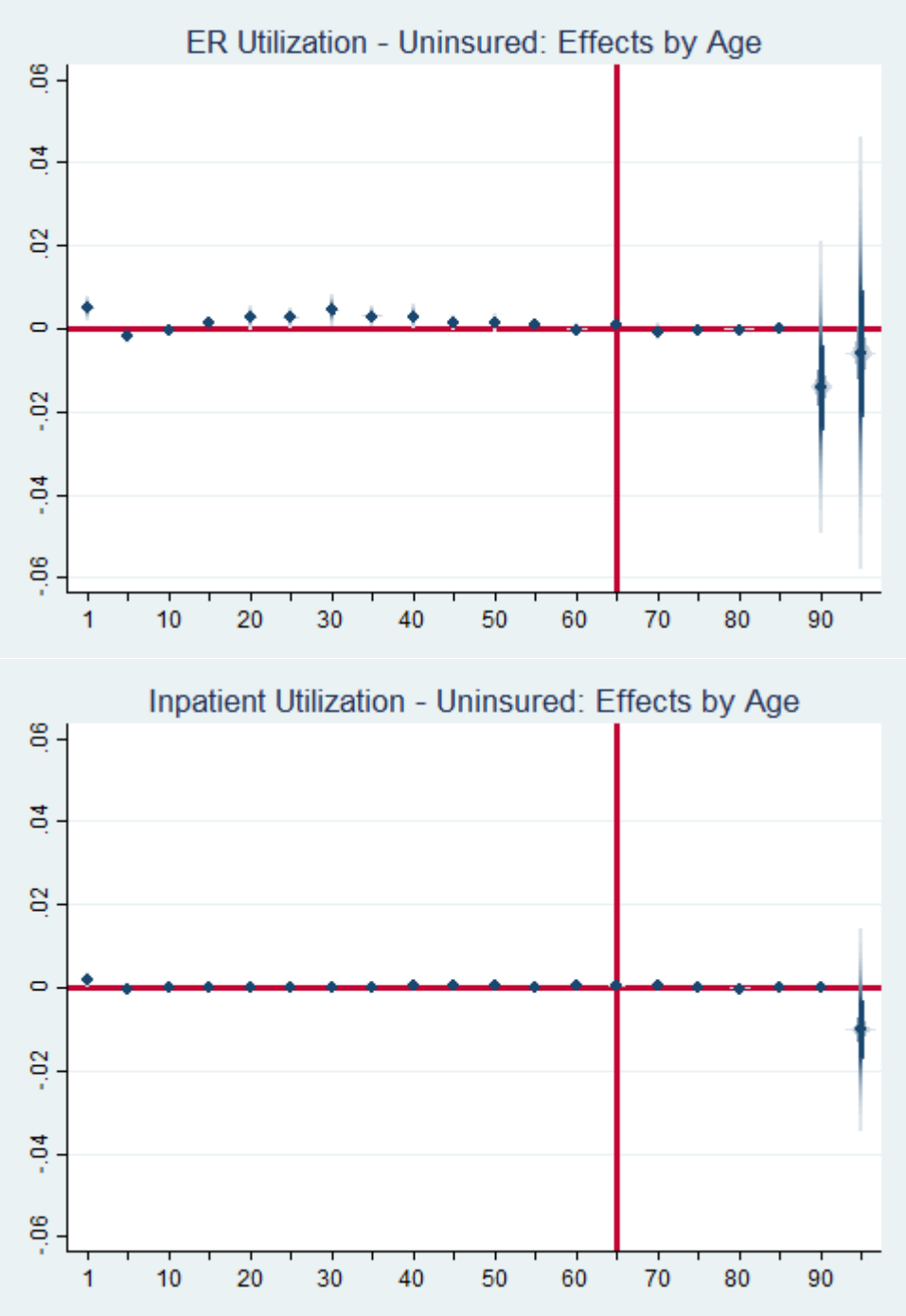


Figure A1: Effects by Age, Inpatient Utilization Funded by Medicaid
Excluding Infants

