The Effect of Taiwan's National Health Insurance on Mortality of the Elderly: Revisited

By

Kang-Hung Chang

China Center for Human Capital and Labor Market Research

Central University of Finance and Economics

November 2009

Abstract

With the help of new data emerged, this paper points out some methodological problems in a recent paper attempting to estimate the effect of Taiwan's National Health Insurance (NHI) on mortality of the elderly but failing to find a significant effect. It is shown that their conclusion can be reversed if the problems are appropriately addressed. Further, this paper adopts an improved method with more recent and detailed data to re-estimate the NHI effect on mortality of the elderly. The main finding suggests that NHI lowered the mortality hazard ratio of the previously uninsured elderly to their continuously insured counterparts by roughly 30%. In addition, the previously uninsured elderly with reported chronic conditions, who arguably had higher mortality risk and needed more medical care than their healthy counterparts, are also found to have experienced a larger NHI effect on both utilization and mortality hazard ratio, suggesting NHI lowered the mortality risk of the elderly through medical care.

Keywords: universal health insurance; mortality; medical care utilization; elderly; Taiwan

JEL codes: I18

1. Introduction

On March 1, 1995, Taiwan initiated a universal health insurance mandate called *National Health Insurance* (NHI), which provided equal benefits to the 21 million citizens on the island, including 8.6 million who were previously uninsured (see Cheng 2003 for a detailed introduction to Taiwan's NHI and insurance market prior to 1995). Its universal and compulsory nature makes Taiwan's NHI a large-scale *natural experiment* rarely seen in recent decades. In particular, it generated arguably exogenous variations in insurance status among the previously uninsured. A growing literature has exploited this feature to estimate the *causal* effects of Taiwan's NHI on medical care utilization and health outcomes (Cheng and Chiang 1997; Chen et al. 2007; Wen et al. 2008). By far, most of the findings in this literature support the view that NHI increased the medical care utilization of the previously uninsured relative to their continuously insured counterparts. Nevertheless, there is no consensus on the NHI effect on health outcomes.

In particular, a recent paper by Chen et al. (2007) adopted a difference-in-differences (DD) method to estimate the NHI effects on the elderly. Their major conclusion is that whereas NHI significantly boosted the medical care utilization of the previously uninsured elderly relative to their continuously insured counterparts, the rise in utilization did not translate into a relative improvement in mortality. They attributed the lack of evidence on mortality to issues such as quality of care, moral hazard, etc.[1] However, I acquire the same data that they used along with a new data set that has not been used in an attempt to replicate their results and find that the lack of evidence on mortality is more likely resulted from their selection of the treatment and control groups as well as some dubious mortality rates that they obtained. I elucidate these two problems

[1] They suggestively attribute the lack of evidence on the mortality effect to four possible reasons: 1) one-year mortality rate may not be a sensitive measure; 2) medical care may not be the major determinant of mortality; 3) there may exist some quality and efficiency problems in the health care delivery system; 4) the increase in utilization may only reflect an increase in moral hazard which had little benefits to health.

in great details in section 2 and 3 respectively and further show that their conclusion could be dramatically reversed if these problems are appropriately addressed.

Moreover, I suggest a new way, which is less error-prone, to select the treatment and control groups. With more recent and detailed data, I incorporate the DD method into a hazard model to re-estimate the NHI effect on mortality. I demonstrate my empirical model and estimation results in section 4. Opposite to what Chen et al. have concluded, I find that NHI lowered the mortality hazard ratio of the previously uninsured elderly to their continuously insured counterparts by roughly 30%.

Conceptually, the NHI effect on mortality is most likely mediated through medical care. In section 5, I estimate the NHI effects on utilization among subgroups of the elderly with chronic conditions, who arguably had higher mortality risk and needed more medical care than their healthy counterparts. I find that the previously uninsured elderly with reported chronic conditions generally experienced a larger relative increase in utilization than their counterparts with no reported chronic conditions. In accordance, these elderly with reported chronic conditions also experienced a larger reduction in mortality hazard ratio. These findings suggest that NHI did improve the health of the elderly through medical care as I conclude in the last section.

2. Selection of the Treatment and Control Groups

Since NHI was a universal program and thus no single Taiwanese citizen was exempt from it, Chen et al. used the elderly with no pre-NHI insurance as the treatment group and the elderly with pre-NHI insurance as the control group. So-called pre-NHI insurance includes several employment-based public programs available in the pre-NHI period such as Government

Employee Insurance (GEI), Labor Insurance (LI), Farmer's Health Insurance (FHI), and Veteran Insurance (VI). Insurees of these public programs were automatically transferred to be covered by NHI after 1995.[2]

The validity of the DD method, which Chen et al. adopted, hinges crucially on the selection of the treatment and control groups (Meyer 1995). In this context, the insurance status of the elderly on the eve of NHI is the key selection criteria but unfortunately unobserved. The longitudinal elderly survey that Chen et al. used—*Surveys of Health and Living Status of the Elderly in Taiwan* (SHLSET)—provided only limited pre-NHI insurance information.[3] They used four waves of SHLSET—1989, 1993, 1996 and 1999. The initial sample of SHLSET was drawn in 1989 and consisted of 4,049 elderly who represented the population aged 60 and older then. Among these four waves, only the 1993 wave asked the elderly about their *current* insurance status.[4] Besides, due to death and other various reasons, only 3,155 of the initial 4,049 elderly were successfully followed up in 1993.[5] Therefore, no pre-NHI insurance information was available for the 894 elderly who were not interviewed in 1993.

In response, they used the 1993 current insurance status to represent the pre-NHI insurance status for the 3,155 elderly.[6] For those not interviewed in 1993, they assigned pre-NHI insurance status to them if they belonged to certain occupation groups in 1989. Specifically, they

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[2] An underlying assumption of using the elderly with pre-NHI insurance as the control group is that the shift from the pre-NHI insurance programs to NHI cast a minimal impact on them at least in terms of medical care utilization and health outcomes.
[3] SHLSET is conducted by Bureau of Health Promotion (formerly known as Taiwan Provincial Family Planning Institute). The initial sample of 4,049 elderly people was drawn in 1989 to represent the population aged 60 and older at the time.
[4] In 1993, the elderly were asked if and what kind of health insurance they currently had. However, the survey did not ask when the elderly started their current insurance. Therefore, their prior insurance status as well as their status between 1993 and 1995 is unknown.
[5] About 66% of the 894 elderly who were not interviewed in 1993 had died before the interview.
[6] In fact, among the 3,155 elderly people in the 1993 survey, 787 and 2,637 elderly people reported being currently uninsured and insured respectively and 1 person did not respond to this question.

assumed one was insured if he or she was a government employee, soldier or farmer in 1989.[7] More importantly, they assumed the insurance status of the elderly remained unchanged from 1989 to 1993 (probably in order to correspond to the 1989 and 1993 waves.) In the end, their analysis sample consisted of 3,899 elderly with 2,990 with pre-NHI insurance and 909 with no pre-NHI insurance, implying an uninsurance rate of 23.3% in pre-NHI period.

While it seems reasonable to assume that the insurance status of the elderly remained unchanged in the pre-NHI period because they tended to begin their employment-based insurance earlier in life, the fact is that a substantial amount of the elderly did change their insurance status in that period. Before NHI was initiated, there was a dramatic, but less noted, decline in uninsurance rate in the late 1980s and early 1990s.[8] I use statistics collected from various administration agencies of the pre-NHI public insurance programs to calculate the uninsurance rate (the proportion of the population not covered by the pre-NHI public programs) from 1989 to 1994. It is worth noting that there was no comprehensive private health insurance at the time that was comparable to the public programs.[9]

Figure 1 illustrates the uninsurance rate for the entire population (solid line) and the aged population of 60 and older (dashed line).[10] As shown, the population uninsurance rate

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[7] They argued that people with these occupations were most likely to be covered by the major public programs at the time. In the 1989 survey, the elderly were asked whether they are currently working, and if yes, what their occupation is. However, only some of the elderly were currently working with known occupation.

[8] The decline was likely resulted from a series of expansions of the public insurance programs. For example, Farmer's Health Insurance started in 1985 in some counties and gradually extended to other counties in the late 1980s and early 1990s. And farmers were a major occupation group of this cohort.

[9] Private insurance then only provided limited benefits and generally did not cover any outpatient service.

[10] The uninsurance rates are my own calculations. The uninsured refer to people not covered by either of Government Employee Insurance, Labor Insurance, Farmers' Health Insurance and Veteran Insurance, which exhaust all available insurance programs at the time. There was no private comprehensive health insurance in Taiwan. Uninsurance rate is defined as the ratio of the uninsured to the population. Uninsurance rates are my own calculations using statistics from the following sources. Population statistics are from *Statistical Yearbook of Interior* published by the Ministry of Interior, which is available at http://www.moi.gov.tw/stat/english/year.asp. Numbers of GEI insurees are from *GEI Statistics* published by Bank of Taiwan (formerly by Bureau of Central Trust), which is available at http://www.bot.com.tw/GESSI/Statics/default.htm. Numbers of FHI and LI insurees are from *Annual Report* published by Bureau of Labor Insurance, which is available at http://www.bli.gov.tw/en/.

significantly dropped from 53% in 1989 to only 40% in 1994. Meanwhile, the elderly uninsurance rate also decreased, although by a smaller magnitude, from 32% in 1989 to 28% in 1990 and to 25% in 1994. This observation suggests that at least some of the elderly in the representative sample of SHLSET should have changed from being uninsured to insured in this period. On the other hand, it is worth noting that the elderly uninsurance rate remained rather stable in the last three years before NHI was implemented—1992, 1993, 1994, suggesting the insurance status of the elderly was less likely to change in these three years.

To further measure how many of the elderly in SHLSET may have changed their insurance status, I acquire the four waves that Chen et al. used along with an additional wave that they did not use.[11] The additional wave is a telephone follow-up conducted in 1991-1992 and focused on only health and life insurance. In this follow-up, the elderly were asked about both their *current* and *past* insurance status, which enable me to compare the insurance status in 1989 (obtained from the telephone follow-up) and 1993 (obtained from the 1993 survey) for those whose insurance status was known in both years.[12] Among the 2,312 elderly—a subset of the 3,155 elderly—whose insurance status in 1989 and 1993 were both known, 408 shifted from being uninsured in 1989 to insured in 1993 and 73 shifted from being insured in 1989 to uninsured in 1993 (Table 1). In other words, about 20% of these elderly had changed their insurance status and the majority of them shifted from being uninsured to insured.

Yet, since we have observed that the elderly uninsurance rate remained quite stable after 1992, the insurance status of the 3,155 was less likely to change after they were interviewed in

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Numbers of VI insurees are kindly provided by Veteran Affairs Commission (VAC). They are not available on-line but can be requested from the VAC.

[11] It is not clear why they did not use it.

[12] In particular, they were asked if and what kind of health insurance they currently had and, if currently insured, when they began the current policy, and, if currently uninsured, whether and when they had insurance before. However, only 3,569 elderly appeared in the follow-up and some of them failed to recall their insurance beginning year. I am only able to learn insurance status in 1989 for 2,312 elderly people whose 1993 status is also known.

1993, suggesting selection of the treatment and control groups based on their 1993 insurance status was credible. The real problem is how to select the treatment and control groups among those not interviewed in 1993. In particular, the downward uninsurance trend suggests that insurance status in 1989 for some elderly, if observed, could have been different from theirs on the eve of NHI.

In addition, assigning pre-NHI insurance to these elderly using their occupations in 1989, as Chen et al. did, suffers several problems. First, the elderly may have changed their occupation or retired after 1989 and thus changed their insurance status. Second, Chen et al. considered neither the private sector workers nor the dependents of government employees and veterans who were respectively eligible for LI, GEI and VI.[13] Third, they claimed that they were able to assign insurance status to 744 elderly people using occupation information in 1989. However, I only found 176 elderly people that were still working and had known occupation in 1989. It is not clear how they assigned insurance status to 744 elderly people. At last, even though their assignments of the 1989 insurance status were correct, they were still subject to changes later.

In sum, while the 1993 insurance status of the 3,155 elderly was likely the same as their unobserved insurance status on the eve of NHI, their assignments of pre-NHI insurance status to the 744 elderly were problematic. The profound implication of this is that Chen et al. could have categorized some elderly people who supposedly should belong to the control group into the treatment group and vice versa, resulting in biased DD estimates of the NHI effects. I show below that this can, at least partially, explain why they did not find any significant NHI effect on mortality.

[13] In principle, among the major public insurance programs, only GEI and VI provided coverage to dependents. However, FHI usually provided coverage to family members of farmers as long as the farm land area per capita in each family met a minimum. This implies that for example, a wife of some farmer might have reported not being working but in fact was covered by FHI. However, using her occupation status would have categorized her as uninsured.

3. DD Estimates of the NHI Effect on Yearly Mortality

It should not be surprising that their selection strategy of the treatment and control groups greatly affected their DD estimates of the NHI effect on yearly mortality.[14] Their yearly mortality rates and the corresponding DD results in their Table III are reproduced in Panel A of Table 2. As shown, the gap in yearly mortality rate between the treatment group (they called it "no pre-NHI insurance group") and the control group (they called it "pre-NHI insurance group") first shrank from 1.54 percentage points in the "before NHI" period (1989 and 1993) to only 0.47 percentage points in 1996 but then widened to 1.21 percentage points in 1999. The two DD estimates suggest the NHI effect were 1.16 and 0.42 percentage points in the case of 1996 and 1999 compared to the "before NHI" period respectively. However, neither was statistically significant.

To see how their selection strategy may have affected their results, I recalculate the mortality rates for the treatment and control group as well as the DD estimates using two mortality data sets—referred as BHP and DOH data—provided by the Bureau of Health Promotion, administrator of SHLSET (Chen et al. should have used the same data sets; see footnote 15 for details about these data).[15] Since I am not able to replicate their assignments of pre-NHI insurance status to the 744 elderly, the best I can do is to calculate the mortality rates of

[14] Chen et al. claimed that they use death certificates of the elderly from 1989 up to November 2000 that come along with SHLSET. In fact, Bureau of Health Promotion (BHP), administrator of SHLSET, did not have death certificates. They collected three sets of death information, including death year and month, in three ways: 1) by interviews (BHP); 2) by linking the elderly to death records at the Department of Health (DOH); 3) by linking the elderly to death records at the Ministry of Interior (MOI). BHP used national ID numbers of the elderly to link them to death records at DOH and MOI. Upon request, BHP provides users of SHLSET these three sets. In other words, unless having acquired death information from other sources, Chen et al. should have these three sets.

[15] There are about 8% mismatches between BHP and DOH data. Among these mismatches, it is difficult to tell which data set is more reliable than the other. Hence, I report results from both sets. Meanwhile, I do not use the set collected from the MOI because it records very few death cases before 1997 (only 3 death cases) due to lack of electronic death records at the MOI in early years.

the treatment and control group using the 3,155 elderly.[16] In fact, this way provides a good way to test how sensitive their results are to the inclusion of the 744 elderly. Moreover, since the 3,155 elderly were still alive at least for some time in 1993, I am not able to calculate the 1989 and 1993 yearly mortality rates. Instead, I use the 1994 yearly mortality rates as the reference. The results are reported in Panel B and C in Table 2.

First note that in column (1), their mortality rate in the 'before NHI' period was 4.97% for the treatment group (Panel A), which is much lower than both the 6.86% (Panel B) and 6.18% (Panel C) in 1994 that I calculate using the BHP and DOH data. This can be resulted either from the inclusion of the 744 elderly or the different reference years. However, in column (4), their mortality rate in the 'before NHI' period for the treatment group was 3.34% (Panel A), which is similar to the 3.73% (Panel B) and 3.52% (Panel C) in 1994 that I calculate. Moreover, in column (2) and (5), their mortality rates in 1996 are also very close to what I obtain. These suggest that the inclusion of the 744 elderly is more likely the main reason that causes the difference in column (1) and that they may have included some "healthy elderly" in the treatment group, who were actually insured in the 'before NHI' period. On the other hand, that the inclusion of the 744 elderly did not cause differences in 1996 and 1999 is likely because the majority of them had died before 1996.[17]

Second, their 1999 rates in column (3) and (6) were both lower than what I obtain. In particular, the 1999 rate for the control group was only 1.87%, which seems puzzling. If NHI did have an impact on mortality, one would expect the big drop in 1999 to have occurred to the treatment group, rather than the control group. More interestingly, this is likely related to not only their selection but also their use of the mortality data. To see this, I calculate the total deaths

[16] In fact, in the 1993 survey, 787 reported they were currently uninsured, 2,367 insured and 1 unknown. I thus discard this case with unknown status.

[17] 70% of them had died before 1996 and 74% of them had died before 1999.

in 1989 plus 1993, 1996 and 1999 implied by their yearly mortality rates and compare them with what I find in the BHP and DOH data.[18] Note that total deaths should have nothing to do with the selection. The total deaths are reported in Table 3. As shown, although Chen et al. used a subsample (n=3,899) smaller than the initial sample (n=4,049), their total deaths in 1989 plus 1993 and in 1996 are reasonably close to the total deaths recorded in the BHP and DOH data. However, their mortality rates imply only 48 total deaths in 1999, less than one third of what the BHP (147 deaths) and DOH (159 deaths) have recorded. This problem is unlikely to be explained by the inclusion of the 744 elderly because more than 70% of them had died before 1999.

By now, it is almost obvious that the inclusion of the 744 elderly and the suspicious few deaths in 1999 are the two main possible factors that make their DD results small and not statistically significant. In their paper, they claimed that they did a sensitivity test by excluding the 744 elderly and also claimed that their results were robust to this test but did not report the test results. However, by using only the 3,155 elderly, my DD results (column (7) and (8) in Panel B and C) suggest that the NHI effect on yearly mortality rate ranged from 2 to 3 percentage points and the results are generally statistically significant, reversing their conclusion.[19]

4. Re-estimate the NHI Effect on Mortality

Previous discussions suggest that the 1993 insurance status of the 3,155 elderly was most likely to be the same as their true unobserved insurance status on the eve of NHI. In this section,

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[18] Note that even though I am not able to replicate their selection the treatment and control group, this does not prevent me from comparing total deaths in the entire elderly sample.

[19] The cluster-robust standard errors are obtained from two simple DD regressions of the yearly mortality outcome variable on a group dummy indicating the treatment group, a year dummy indicating the year of 1996/ 1999 and an interaction term of the two dummies.

I thus use only these 3,155 elderly with a mortality hazard model to re-estimate the NHI effect on mortality while controlling for their observed baseline characteristics. In addition, since I have more recent and detailed monthly mortality data up to the end of 2003, my analysis period spans from March 1993 to the end of 2003. To clean the data, I discard one elderly with unknown insurance in 1993, ten who reported being covered by private insurance and two who reported being insured by unknown insurance plans.[20] The final analytic sample consists of 3,136 with 785 uninsured (the treatment group) and 2,351 insured (the control group) in 1993. Table 4 summarizes their baseline characteristics in 1993. As shown, the treatment group was more likely to be women, less educated, poorer, and less healthy.

I incorporate the DD strategy into an exponential hazard model described below.

$$(1)\quad h(t_i) = \exp(\beta_0 + \beta_1 TREAT_i + \beta_2 POST_i(t_i) + \beta_3 TREAT_i \times POST_i(t_i) + X_i B),$$

where h is the exponential hazard function; t_i is the number of months that individual i was alive during the period from March 1993 to the end of 2003 and assumed to be exponentially distributed; *TREAT* is an indicator for the treatment group; *POST* is an indicator for the post-NHI period and dependent on one's survival time, t; *POST* equals one if one was still alive in the post-NHI period and zero otherwise; X is a vector of baseline controls evaluated in 1993 including age, sex, education, ethnicity, marital status, residence region, living arrangement, employment status, monthly income, activities of daily living (ADL), chronic conditions, self-reported health and three health behavior measures. Based on the mortality hazard function, a likelihood function is constructed accordingly and the model is then estimated by maximum likelihood method. The BHP and the DOH data sets are used in separate estimations.

[20] The reason is that private insurance in Taiwan only provided limited inpatient benefits. Privately insured elderly were thus not comparable to those covered by comprehensive social insurance programs. Neither were they comparable to those completely uninsured.

If we take the natural logarithm on both sides of equation (1), β_3 measures the difference in differences in log mortality hazards conditional on baseline characteristics and is the DD estimator of the NHI effect. Alternatively, it is easier to interpret the estimation results in terms of mortality hazard ratio. In equation (1), exp (β_1) is the mortality hazard ratio of the treatment group to the control group in the pre-NHI period, while exp $(\beta_1 + \beta_3)$ is the mortality hazard ratio in the post-NHI period. If NHI lowered the mortality hazard ratio, it is expected that exp $(\beta_1) >$ exp $(\beta_1 + \beta_3)$.

It is worth noting that this model is a *piecewise constant hazard model* because the baseline mortality hazard, $\exp(\beta_0 + \beta_2 POST_i(t_i))$, which is common to both the treatment and control group, remains constant within each period and faces a possible discrete jump at the point when NHI was initiated if β_2 is positive. Assuming a constant hazard in each period may not be sensible for the case of elderly. However, for my DD strategy, the exact shape of the baseline hazard function is not the main concern.

Estimation results are reported in column (1) and (2) in Table 5. As shown, the estimate of the NHI effect, β_3, is -0.25 for both the BHP and the DOH data and both are statistically significant at 10% level, suggesting the NHI reduced the log mortality hazard for the treatment group relative to the control group. Further, the DD estimates imply the mortality hazard ratio of the treatment group to the control group dropped from 1.38 (1.41) in the pre-NHI period to 1.07 (1.1) in the post-NHI period for the BHP (DOH) data. This suggests that after the implementation of NHI, the mortality hazard ratio had dropped by about 30%.

Furthermore, to check if my results are sensitive to distributional assumption, I also estimate a similar model which assumes that the survival time, *t*, has a Weibull distribution and allows the baseline hazard to increase as the elderly age.[21] The results are reported in column (3) and (4) in Table 5. As shown, the results are similar.

5. The NHI Effects on Utilization and Mortality of the Elderly with Chronic Conditions

It is of interest to further ask how NHI had led to the drop in mortality hazard ratio. In principle, it is most plausible that the effect was mediated through medical care utilization. Especially, one particular group of the elderly who arguably had higher mortality risk and needed more medical care than others is the elderly with chronic conditions. It is thus a reasonable conjecture that if NHI really helped, it is the previously uninsured elderly with chronic conditions who benefited more than others through medical care, which eventually led to a lower mortality risk. Table 6 summarizes the prevalence of common chronic conditions reported by the elderly in SHLSET. As shown, the majority of the elderly (about 75%) had reported at least one condition and the average was about 1.6 conditions. The most prevalent conditions include heart problems, hypertension, cataract, diabetes, upper respiratory problems, etc. However, the distributions of these conditions in these two groups are similar.

To see if NHI had a larger effect on utilization for the elderly with chronic conditions, I use a linear DD model to estimate the NHI effects among the subgroups with chronic conditions. The model is described as follows.

$$(2)\quad Y_{it} = \beta_0 + \beta_1 TREAT_i + \beta_2 YR96_t + \beta_3 TREAT_i \times YR96_t + X_i B + \varepsilon_{it},$$

where Y_{it} is the utilization outcome for individual i at time t; $YR96$ is a year dummy for 1996; other notations are the same as in equation (1). Again, β_3 is the DD estimator. I use two years of

[21] For Weibull, the hazard function is $h(t_i) = pt^{p-1}\exp\left(\beta_0 + \beta_1 TREAT_i + \beta_2 POST_i(t_i) + \beta_3 TREAT_i \times POST_i(t_i) + X_i B\right)$ where p is the so-called "shape" parameter. If p>1, the baseline hazard monotonically increases over time. In my estimation, p is about 1.4. However, the increasing baseline hazard does not affect the mortality hazard ratio between the two groups because it is common to both groups.

data—1993 and 1996—and keep only those appear in both waves with no missing values in order to avoid a potential nonrandom attrition bias due to death. Outcome variables include a dummy variable indicating if one had any outpatient visits in the past month and a dummy variable indicating if one had ever been hospitalized in the last year. Regressions are run by chronic condition groups separately.

For brevity, only the interaction term in equation (2) is reported in Table 7. As shown, compared to elderly with no reported chronic conditions, those with any reported chronic conditions generally experienced larger NHI effects on the probability of outpatient visits and the probability of hospitalizations.[22] In other words, a previously uninsured elderly with certain chronic conditions generally had a larger *relative* increase in utilization due to NHI than their counterparts without chronic conditions. Moreover, the effects were particularly large for the previously uninsured elderly with kidney problems, hypertension and diabetes. This should not be surprising because these conditions generally require intensive uses of medical care either in the form of outpatient visits or hospitalizations.

Following the findings in utilization, I estimate equation (1) again using these subgroups to see if the relative large increase in utilization also translated into a relative large reduction in mortality hazard ratio. For brevity, only the mortality hazard ratio is reported in Table 8. As shown, compared to the elderly with no chronic conditions, the elderly with chronic conditions in general did also experience a larger reduction in mortality hazard ratio. These findings suggest that the NHI lowered the mortality hazard ratio most likely by improving utilization of medical care to the previously uninsured elderly. The effects were particularly large among the elderly with chronic conditions.

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[22] Note that here the elderly "with no chronic conditions" refer the elderly who did not report any chronic conditions listed in the survey. It is possible that they had other chronic conditions not listed in the survey.

6. Conclusions

The study of the NHI effects in Taiwan using the DD method with SHESLT faces an obstacle of unobserved insurance status on the eve of NHI, which is crucial in determining the treatment and control groups. With the help of new data emerged, I show that the main reasons that Chen et al. failed to find a significant NHI effect on mortality are likely their selection of the treatment and control groups as well as their suspicious mortality rates in 1999.

One simple way to improve this research is to use only those elderly with known 1993 insurance status because their 1993 status was most likely to be the same as their unobserved insurance status on the eve of NHI. Hence, selection of the treatment and control groups based on their 1993 status is most credible. Following this strategy and using more recent and detailed data, my re-estimation results of the NHI effect on mortality have overturned the conclusion by Chen et al. In particular, I find that NHI had lowered the mortality hazard ratio of the previously uninsured elderly to their continuously insured counterparts by about 30%. I further find that the effect was particularly large among the elderly with reported chronic conditions, who arguably had higher mortality risk and needed more medical care than those with no reported chronic conditions.

It is interesting to point out that I also find similar NHI effects, although not reported here, on utilization to what Chen et al. reported in their paper—NHI greatly increased medical care utilization for the previously uninsured elderly relative to their continuously insured counterparts. The main reason for the similar findings is that utilization information was only available since the 1993 wave of SHLSET. Therefore, their sample for the estimation of the NHI effect on utilization consisted of only those elderly appeared in 1993, the same group that I also use.

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Figure 1. Uninsurance Trend 1989-1994

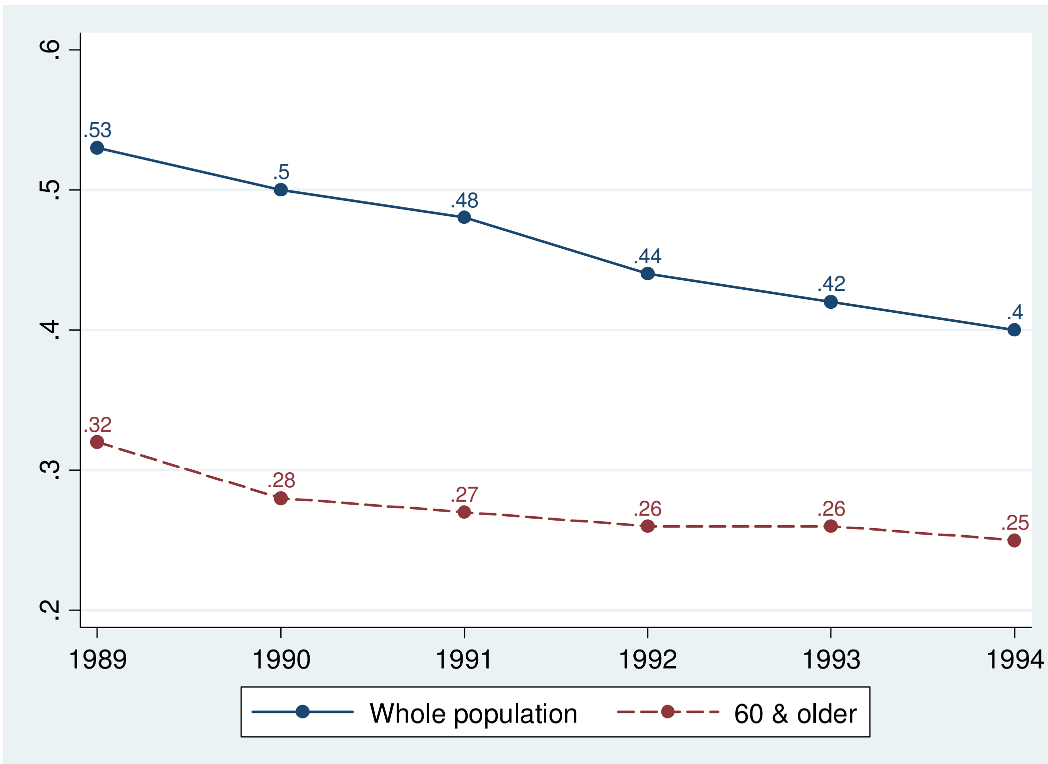

Table 1. Comparison of Insurance Status of the Elderly in 1989 and 1993

	1993 Status		Total
1989 Status	Insured	Uninsured	
Insured	1,520	73	1593
Uninsured	408	311	719
	1,928	384	2,312

Notes: insurance status in 1989 is inferred from the 1991-1992 telephone follow-up; only the elderly who were interviewed in both the 1991-1992 telephone follow-up and the 1993 wave with known insurance status in both years are included in this table.

Table 2. Comparison of One-year Mortality Rates and DD Results (%)							
'No pre-NHI insurance' (Treatment Group) (1) ~ (3)			'Pre-NHI insurance' (Control Group) (4) ~ (6)			Difference-in-differences (7) ~ (8)	
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A. Chen et al.							
Before NHI	1996	1999	Before NHI	1996	1999	DD1	DD2
4.97	4.85	3.08	3.34	4.38	1.87	-1.16	-0.42
(1,490)	(515)	(455)	(5,509)	(2,125)	(1,823)	[-1.1]	[-0.19]
Panel B. Use BHP Data							
1994	1996	1999	1994	1996	1999	DD1	DD2
6.86	4.85	5.5	3.73	4.56	5.77	-2.84	-3.4
(758)	(660)	(564)	(2,303)	(2,125)	(1,820)	[-2.06]	[-2.28]
Panel C. Use DOH Data							
1994	1996	1999	1994	1996	1999	DD1	DD2
6.18	4.8	5.93	3.52	4.26	6.11	-2.12	-2.85
(760)	(666)	(573)	(2,304)	(2,137)	(1,848)	[-1.59]	[-1.92]

Notes: DD1 = ((2)-(1))-((5)-(4)); DD2=((3)-(1))-((6)-(4)); observations are in parentheses; *z*-scores are in brackets for Chen et al. DD results; cluster-robust *t* statistics are in brackets for DD results in panel B and C; results of Chen et al. are directly adopted from their Table III; one-year mortality rates for BHP and DOH are based on the 3,155 elderly whose 1993 insurance status were known; cluster-robust *t* statistics are calculated based on a simple DD regression of a death indicator on a group dummy indicating the treatment group, two year dummies for 1996 and 1999 and two interaction terms of the group dummy with the year dummies; "before NHI" period refers to 1989 and 1993 in Chen et al.

Table 3. Comparison of Total Deaths			
	1989+1993	1996	1999
Chen et al. (n=3,899)	258	118	48
BHP (n=4,049)	248	140	147
DOH (n=4,049)	236	132	159

Notes: total deaths in the first row are calculated based on the mortality rates and numbers of observations in Panel A in Table 2.

Table 4. Characteristics of the Elderly in 1993		
	(1) Control (Insured)	(2) Treatment (Uninsured)
Age (mean)	71	72
Female (%)	39	57
Married with spouse present (%)	67	47
Ethnicity (%)		
Minnan	56	74
Hakka	17	10
Mainlander	24	14
Aboriginal	2	1
Region (%)		
East	9	6
North	23	38
Central	36	29
South	32	26
Urban area (%)	59	77
Live with adult children (%)	64	70
Years of education (%)		
0 Year	45	60
1-6 Years	33	31
7 Years and more	22	9
Currently employed/ self-employed (%)	23	11
Monthly income > NT$10,000 (%)	45	31
Self-reported health (%)		
Very good/ good	41	36
Fair	33	31
Bad/ very bad	21	24
No response	5	9
Functional limitation (ADL) (mean)	0.38	0.54
Have one or more chronic condition(s) (%)	75	74
Health behavior (%)		
Smoking	29	27
Drinking	13	13
Betel nut chewing	6	4
Total	2,351	785

Notes: ADL is an average score of 12 functional limitation items of daily activities; each item ranges from 0 (no difficulty) to 3 (cannot do it at all); chronic conditions include arthritis, upper respiratory problems, cancer, cataract, diabetes, hypertension, heart problems, kidney problems, liver problems, and stroke.

Table 5. Re-estimation of the NHI Effect on Mortality				
	(1)	(2)	(3)	(4)
	Exponential w/ BHP Data	Exponential w/ DOH Data	Weibull w/ BHP Data	Weibull w/ DOH Data
TREAT	0.32	0.35	0.33	0.35
	(0.13)	(0.14)	(0.13)	(0.14)
TREAT×POST	-0.25	-0.25	-0.25	-0.25
	(0.15)	(0.15)	(0.15)	(0.15)
Subjects	3,136	3,136	3,136	3,136
Deaths	1,367	1,389	1,167	1,389
Mortality Hazard Ratio:				
Pre-NHI	1.38	1.41	1.38	1.42
Post-NHI	1.07	1.1	1.08	1.11

Notes: analysis period starts from March 1993 to the end of 2003; standard errors are in parentheses; baseline control variables include age, sex, ethnicity, living region, urban area, living arrangement, education, marital status, monthly income, chronic conditions, ADL, self-reported health, smoking, drinking and chewing betel nuts; estimates of the constant term, POST and controls are not reported for brevity; pre-NHI mortality hazard ratio=exp(TREAT); post-NHI mortality hazard ratio=exp(TREAT+TREAT×POST)

Table 6. Prevalence of Reported Chronic Conditions (%)

Conditions (%)	(1) Treatment	(2) Control
Arthritis	26	24
Upper respiratory problems	14	17
Cancer	2	2
Cataract	27	25
Diabetes	11	10
Heart problems	21	21
Hypertension	28	31
Kidney problems	9	6
Liver problems	6	6
Stroke	7	7
No above conditions	26	25
Average conditions	1.6	1.6
n	785	2,351

Table 7. Estimation of the NHI Effect on Utilization by Subgroups with Reported Chronic Conditions

	(1) Outpatient	(2) Hospitalization
(A) With Any Conditions	0.145	0.107
	(0.035)	(0.028)
	[1,830]	[1,836]
(B) With Diabetes	0.209	0.086
	(0.09)	(0.086)
	[235]	[236]
(C) With Hypertension	0.106	0.111
	(0.055)	(0.049)
	[751]	[754]
(D) With Kidney Problems	0.224	0.215
	(0.113)	(0.088)
	[163]	[164]
(E) No Reported Conditions	0.092	0.060
	(0.054)	(0.036)
	[674]	[676]

Notes: sample consists only those appeared in both 1993 and 1996 to avoid a potential attrition bias due to death; possible chronic conditions include arthritis, upper respiratory problems, cancer, cataract, diabetes, heart problems, hypertension, kidney problems, liver problems, stroke; "no conditions" means have no conditions listed here; elderly in row (B) to (D) may also have other conditions not listed in SHLSET; cluster-robust standard errors in parentheses; observations in brackets; only the interaction terms are reported.

Table 8. Estimation of the NHI Effect on Mortality Hazard Ratio By Subgroups with Reported Chronic Conditions

	(1) Exponential w/ BHP Data	(2) Exponential w/ DOH Data
(A) With Any Conditions		
Pre-NHI	1.35	1.43
Post-NHI	1.05	1.1
(B) With Diabetes		
Pre-NHI	1.86	1.76
Post-NHI	1	1.05
(C) With Hypertension		
Pre-NHI	2.05	2.06
Post-NHI	0.97	0.84
(D) With Kidney Problems		
Pre-NHI	1.67	2.17
Post-NHI	0.49	0.64
(E) No Reported Conditions		
Pre-NHI	1.57	1.35
Post-NHI	1.13	1.15

Notes: control variables include age, sex, ethnicity, living region, urban area, living arrangement, education, marital status, monthly income, chronic conditions, daily activity measure (ADL), self-reported health, smoking, drinking and chewing betel nuts; only hazard ratios are reported; pre-NHI mortality hazard ratio=exp(TREAT); post-NHI mortality hazard ratio=exp(TREAT+TREAT×POST); elderly in row (B) to (D) may also have other conditions not listed in SHLSET.