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

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Abstract

While it is well-accepted that expanding insurance coverage leads to greater health care utilization, it is less clear that health improves as a result. A recent paper by Chen et al. (2007) examined the effect of Taiwan's National Health Insurance (NHI) on mortality of the elderly by comparing mortality of the previously-uninsured and continuously-insured before and after NHI was adopted but did not find a significant effect. I use the same data source and a similar methodology but reach a different conclusion—NHI lowered the mortality hazard of the previously-uninsured elderly relative to their continuously-insured counterparts by roughly 30%. Differences between us stem from different choices in defining the "treatment" and "control" groups, differences in interpretation of the mortality data, and some other modeling decisions. I also find that the previously-uninsured elderly who reported chronic conditions generally experienced larger NHI effects on both utilization and mortality than those without such conditions.

Keywords: universal health insurance; mortality; 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 (Cheng 2003).[1] The universal and compulsory nature of NHI 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 medical care utilization of the previously uninsured relative to the continuously insured—those insured both before and after NHI was adopted. 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 largely increased the medical care utilization of the previously uninsured elderly (the treatment group) relative to their continuously insured counterparts (the control group), the rise in utilization did not translate into a significant improvement in mortality. They attributed the lack of evidence on mortality to low quality of care, moral hazard, etc.[2] However, I acquire the same data that they used 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 their interpretation of the mortality data. I elucidate these two issues in details in section 2 and 3.

[1] See Cheng (2003) for a detailed introduction to NHI and prior insurance programs in Taiwan.

[2] They 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.

Moreover, I adopt an arguably less error-prone way to redefine the treatment and control groups. I then incorporate the DD method into a hazard model to re-estimate the NHI effect on mortality with more recent and detailed mortality data. 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 of the previously uninsured elderly relative 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 subgroups of the elderly with reported chronic conditions, which were prevalent among the elderly. These elderly groups 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 larger NHI effects on both utilization and mortality than their counterparts with no such conditions. This finding is consistent with the experience from the RAND Health Insurance Experiment in the U.S. in the 1970s (Manning et al. 1987).

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.[3] Then they adopted a DD method to estimate the NHI effects.

The validity of the DD method 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 in the data. They used four waves—1989, 1993, 1996 and 1999—of a longitudinal elderly data set called *Surveys of Health and Living Status of the Elderly in Taiwan* (SHLSET). The initial sample of SHLSET was drawn in 1989 and consisted of 4,049 elderly who represented the population aged at least 60. Among these four waves, only the 1993 wave asked the elderly about their *current* health insurance status.[4] Besides, due to death and other 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 assumed one was insured if he or she was a government employee, soldier or farmer in 1989 and uninsured otherwise.[7] More importantly, they assumed that their insurance status remained unchanged from 1989 to 1993 (probably in order to correspond to the 1989 and 1993 waves). In the end, their

[3] 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. These programs provide similar and comprehensive medical benefits (Cheng and Chiang 1997).
[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.
[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.

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 their intent to include the whole initial sample in their analysis is admirable, imputing insurance status in 1989 based on occupation information probably had introduced too many measurement errors (see Appendix 1). On the other hand, even if their imputation was correct, the assumption that the 1989 and the 1993 status were identical was against the fact. Before NHI was initiated, there was a dramatic decline in uninsurance rate in the late 1980s and early 1990s.[8] Figure 1 illustrates the uninsurance rate for the entire population (solid line) and the aged population of 60 and older (dashed line) from 1989 to 1994.[9] As shown, the population uninsurance rate significantly dropped from 53% in 1989 to only 40% in 1994. Meanwhile, the elderly uninsurance rate also decreased 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.

It is worth noting that the elderly uninsurance rate remained rather stable in the last three years before NHI was implemented, suggesting the 1993 insurance status is probably the best pre-NHI insurance information.

[8] The decline was 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] 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 at the time. Uninsurance rate is defined as the ratio of the uninsured to the entire population. Statistics are 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/. 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.

Further, I compare insurance status in 1989 and 1993 for a subset of the elderly in SHLSET. Their 1989 status is learned from a telephone follow-up conducted in 1991-1992 that Chen et al. did not use. [10] In this follow-up, the elderly were asked about both their *current* and *past* insurance status, which allows me to determine the insurance status in 1989 at least for those who appeared in the follow-up.[11] Among the 2,312 elderly whose insurance status in 1989 and in 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 this elderly subgroup had changed their insurance status and the majority of them shifted from being uninsured to insured.

The profound implication of above discussions is that by including the elderly with imputed insurance status in 1989, Chen et al. likely had categorized some elderly people already insured in the pre-NHI period into the treatment group and vice versa. Since most of them changed from being uninsured to insured, the estimated NHI effects were most likely biased towards zero. I show this in the following section.

3. DD Estimates of the NHI Effect on Yearly Mortality

Based on their definition of the treatment and control groups, Chen et al. adopted the DD method to estimate the NHI effect on yearly mortality.[12] Their yearly mortality rates and DD

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[10] It is not clear why they did not use it.
[11] 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.
[12] Chen et al. claimed that they used death certificates of the elderly from 1989 up to November 2000 that came 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

results (originally reported in Table III in their paper) 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 decreased 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. These results seem to show some NHI effect on mortality but neither was statistically significant.

I calculate the mortality rates and the DD estimates using two mortality data sets—referred as BHP and DOH data—provided by the Bureau of Health Promotion, administrator of SHLSET.[13] I only use the 3,155 elderly who had the best pre-NHI insurance information.[14] This provides a good way to test how sensitive their results are to the inclusion of the elderly with imputed insurance status. 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 to represent the "before NHI" period mortality rates. 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 elderly with imputed insurance status or the different reference

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to death records at DOH and MOI. Upon request, BHP provided 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.

[13] 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.

[14] 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.

year. 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—when most of the elderly with imputed insurance status had died—are also very close to what I obtain.[15] These suggest that the inclusion of the elderly with imputed insurance status is more likely the main reason that caused the difference in column (1). In particular, they likely had included some “healthy” elderly in the treatment group, who were actually insured at least for some time in the ‘before NHI’ period.

Second, their 1999 rates in column (3) and (6) were both lower than what I obtain. Especially, their 1999 mortality 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. Moreover, this unexpected drop is unlikely to be caused by the inclusion of the elderly with imputed insurance status because most of them (about 70%) had died by 1999. Instead, it is probably related to their “interpretation” of the mortality data. To see this, I calculate the total deaths 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.[16] 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.

[15] 70% of them had died before 1996 and 74% of them had died before 1999.
[16] 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.

In their paper, they claimed that they did a sensitivity test by excluding the elderly with imputed insurance status and found similar results, which were not reported in their paper. However, my results, which are based on only the 3,155 elderly and supposed to be similar to their sensitivity test, suggest that the NHI effect on yearly mortality rate ranged from 2 to 3 percentage points and the results are generally statistically significant (the last two columns in Panel B and C).[17]

4. Re-estimate the NHI Effect on Mortality

The linear DD estimates in Table 2 have two additional problems. First, they have not controlled for observed differences between the two groups that are potentially related with mortality.[18] It will be shown later that the previously uninsured elderly were more likely to be women, less educated, poorer and less healthy than the continuously insured elderly. Second, the linear DD model fails to fully utilize mortality information at time points other than 1989, 1993, 1996 and 1999.

To improve the estimation, I use the 3,155 elderly, who had the best pre-NHI insurance information, to avoid the measurement errors resulted from imputing insurance status. I incorporate the DD method into a mortality hazard model with observed baseline controls. Instead of using only few yearly data points, I use monthly mortality data from March 1993 to the end of 2003. My analysis period has to begin in March 1993 because the 3,155 were still alive then.

[17] 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.

[18] Chen et al. also estimated a probit model with baseline controls. However, their estimated NHI effect on mortality was still not significant. Again, the key issues are still the selection of the treatment and control groups as well as their interpretation of the mortality data. And the probit model does not address these issues.

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.[19] 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 mortality 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 depends on one's survival time; 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.

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. However, it is difficult to interpret the numerical meaning of the

[19] 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.

coefficient. 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.

It is worth noting that this model is a *piecewise constant hazard model* because the baseline mortality hazard, $\exp(\beta_0 + \beta_2 POST_i(t_i))$, remains constant within each period and faces a possible discrete jump at the point when NHI was initiated. 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. It will be shown soon that the distributional assumption does not affect the DD estimates.

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, 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.[20] The results are reported in column (3) and (4) in Table 5. As shown, the results are almost identical.

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

[20] For Weibull, the hazard function is $h(t_i) = pt^{p-1}\exp(\beta_0 + \beta_1 TREAT_i + \beta_2 POST_i(t_i) + \beta_3 TREAT_i \times POST_i(t_i) + X_i B)$ where 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.

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 such as diabetes, heart problems, hypertension, etc. It is thus a reasonable conjecture that if NHI really helped, the previously uninsured elderly with such chronic conditions would benefit more than others.

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.

$$(2)\quad Y_{iy} = \beta_0 + \beta_1 TREAT_i + \beta_2 YR96_y + \beta_3 TREAT_i \times YR96_y + X_i B + \varepsilon_{iy},$$

where Y_{iy} is the utilization outcome for individual i in year y; $YR96$ is a year dummy for 1996; other notations are the same as in equation (1). Again, β_3 is the DD estimator of the NHI effect.

I use two years of 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.[21] 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.

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. As shown, the reasons that Chen et al. did not find a significant NHI effect on mortality are likely the inclusion of the elderly with imputed insurance status and their interpretation of the mortality data.

[21] 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.

One simple way to improve this research is to use only those elderly with the best pre-NHI insurance information, that is, those elderly appeared in 1993. Selection of the treatment and control groups based on their 1993 status is most credible. My re-estimation results of the NHI effect on mortality show that NHI lowered the mortality hazard of the previously uninsured elderly relative 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. This finding is consistent with the experience from the RAND Health Insurance Experiment in the U.S. in the 1970s.

It is also worth pointing out that the DD method in this case relies on a strong assumption that the continuously insured elderly is an appropriate “control” group in that they can be used to measure the counterfactual time trend in mortality that would have been experienced by the treatment group in the absence of NHI. This assumption may not hold, since we have known that the treatment group was more likely to be women and less healthy than the control group. Yet, the two effects move in opposite directions. On one hand, the mortality rate of the treatment group would increase slower than the control group, because women were known to have a longer life expectancy than men in Taiwan. On the other hand, the mortality rate of the treatment group would accelerate faster than the control group, simply because they were sicker at the baseline. In the end, the net effect depends on which one dominates. However, it is safe to argue that the potential bias in estimating the time trend should be of less concern, since the two effects offset each other.

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Appendix 1

Imputing pre-NHI insurance status to the elderly using their occupations in 1989 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.[22] 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 imputations of the 1989 insurance status were correct, they were still subject to changes later.

[22] 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.

Figure 1. Uninsurance Trend 1989-1994

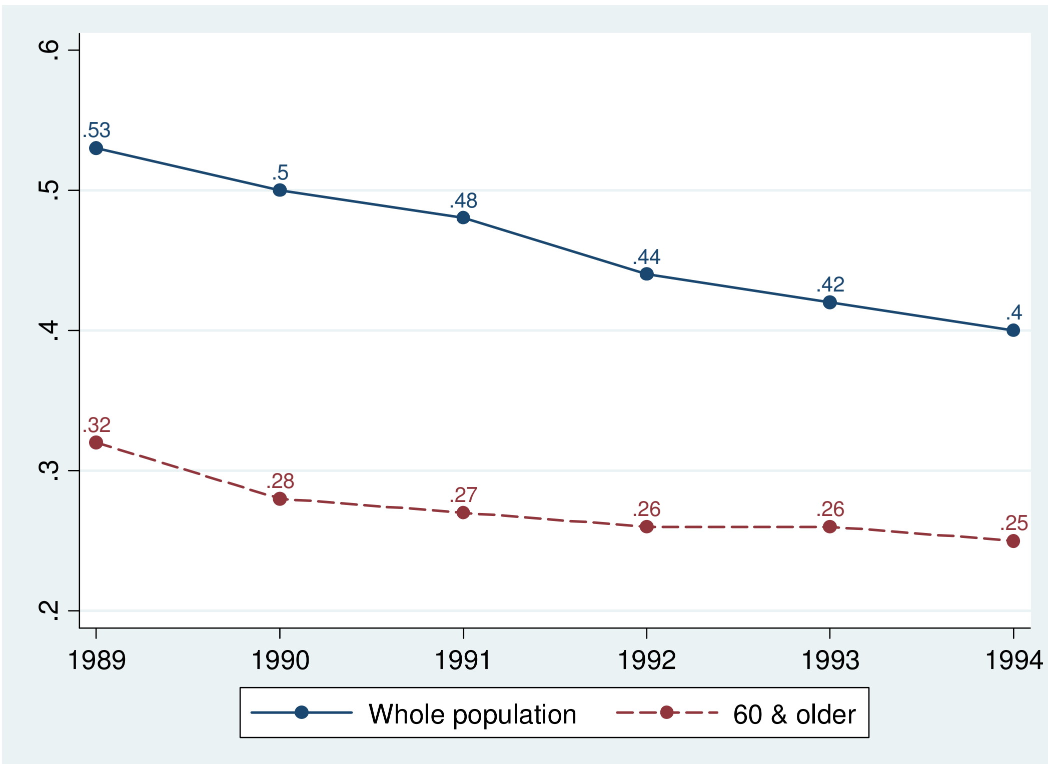

Table 1. Comparison of Insurance Status of the Elderly in 1989 and 1993			
	1993 Status		
1989 Status	Insured	Uninsured	Total
Insured	1,520	73	1593
Uninsured	408	311	719
	1,928	384	2,312

Notes: insurance status in 1989 is learned from the 1991-1992 telephone follow-up; insurance status in 1993 is learned from the 1993 wave; only the elderly whose insurance status was known 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 in Panel A; cluster-robust *t* statistics are in brackets for DD results in panel B and C; results of Chen et al. are adopted from their Table III; one-year mortality rates in Panel B and C only include the elderly who appeared in the 1993 survey; 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 for Chen et al. 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) Exponential w/ BHP Data	(2) Exponential w/ DOH Data	(3) Weibull w/ BHP Data	(4) 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 (%)

	(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 (E) may 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 (E) may have other conditions not listed in SHLSET.