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Channels of Stabilization in a System of Local Public Health Insurance: The Case of the National Health Insurance in Japan*

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Abstract

There are more than 1,700 municipalities serving as insurers in Japan’s system of National Health Insurance (NHI). The NHI has several institutional routes to buffer local premiums from abrupt changes in regional health demands that destabilize the NHI benefit expenditures. After briefly introducing the system of public health care in Japan, this study elaborates on the methods for quantifying the degree of stabilization of local public health care expenditures by critically evaluating the methods that have been utilized in the related literature and proposes a modified method appropriate for this study. It then quantifies the channels and degrees of stabilization using municipal NHI data in the 2000s.

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1. Introduction

The system of public health insurance is broadly classified into two groups: single-payer and multi-payer systems (Hussey and Anderson 2003). The Japanese system of public health insurance may be categorized as a non-competitive multi-payer system with two schemes of public health insurance: employment-based and residence-based. The National Health Insurance (NHI) is a residence-based system of public health insurance where municipalities insure those who are excluded from the employment-based scheme. This regional scheme may be comparable to the pre-1992 regional sickness funds in Germany and to the pre-2000 health insurance societies for the self-employed in Korea, the latter of which Hwang (2008) categorized as following the “traditional” German model. However, Germany made its system “competitive” in 1992 so that the insured can choose among different sickness funds, and which reduced the number of regional funds from 269 in 1993 to only 17 in 1999. Meanwhile, Korea transformed its multi-payer system into a single-payer system in 1999. Therefore, Japan’s NHI is arguably the only non-competitive multiple payer system based on the now-defunct traditional German model.

It is argued that a multi-payer system has its own disadvantages over a single-payer system. For example, Hussey and Anderson (2003) argued that it would be less effective in collecting revenues, controlling costs, and subsidizing health care for low-income individuals. More importantly, it is less effective in pooling risks. The law of large numbers dictates that risks, albeit unpredictable at the individual level, become more predictable as the group size becomes larger (Boadway and Bruce 1980). Since the NHI currently consists of more than 1,700 municipalities (i.e., insurers) that have small enrollments on average, it could not effectively spread health risks across the insured without any inter-municipal transfers. For example, the smallest enrollment was 92 in fiscal year (FY) 2009. In addition, the municipal NHI enrolls a riskier population than that in the other employment-based scheme. Such a riskier population includes self-employed non-professionals, the retired, and the unemployed. To counteract the disadvantages of the NHI system, therefore, there are both a system of inter-institutional transfers from the employment-based public health schemes and a system of inter-regional transfers within the NHI system that involves multiple layers of the government. Such fiscal transfers are expected to smooth changes in municipal NHI premiums imposed on the insured local population—changes caused by shocks to local health demand.

The purpose of this study is to measure the degree of such smoothing effects of the transfers on the nexus between the changes in the NHI medical benefits (which reflect regional health care demand) and changes in the NHI premiums (which reflect fiscal burdens on the insured in a particular municipality). In particular, I examine how the volatility in regional health care demand is tamed through fiscal transfers and cost-sharing schemes at different levels of the government. Meanwhile, it is possible to argue that volatile municipal health demand may not be a serious issue even without the fiscal transfers, since municipalities as NHI insurers can borrow and lend to offset such adverse effects. To see how this argument applies to the Japanese case, I also examine the extent to which such inter-temporal adjustments contribute to smoothing changes in the NHI premiums.

The smoothing effects of interregional fiscal transfers has been a key focus of empirical studies of fiscal federalism (Bayoumi and Masson 1995; Asdrubali et al. 1996; Doi 2000; Buettner 2002; Decressin 2002; Mélitz and Zumer 2002; Andersson 2004; Jüßen 2006; Andersson 2004, 2008; Ramos and Coimbra 2009; Arachi 2010; Furceri 2010; Hepp and von Hagen 2011; Balli et al. 2011). The current study distinguishes itself from the previous studies and contributes to the literature in both substantive and methodological respects as follows. First, while the literature concerns the nexus between gross income and disposal income (or consumption), I focus on the relations between health care demands and premium collections. In this study, the "risk variable" to which shocks occur is the NHI medical benefits, and the "target variable" to be smoothed is the NHI premiums. Therefore, "stabilization" or "smoothing" here involves providing localities with more stable flows of revenues, and preventing large changes in premium collections imposed on local residents. To the best of my knowledge, no study has examined the role of fiscal transfers for locally managed public health insurance as such, nor quantified their stabilizing effects within a multi-payer public health insurance system.

Second, I elaborate on the relation between two strands in the previous studies that quantify the degree of stabilization effects, and relate them to the method of this study. I start with the method by Asdrubali et al. (1996), which has been applied in a number of studies (Doi 2000; Buettner 2002; Andersson 2004; Jüßen 2006; Ramos and Coimbra 2009; Furceri 2010; Balli et al. 2011). This method, which uses two types of identity, decomposes the variance of a risk variable into its sub-factors that measure their respective contributions in buffering the target variable from the effects of the volatile risk variable. I then elaborate on the other strand in the literature. Such studies, in a somewhat ad-hoc manner, regressed the target variable on the risk variable to obtain a measure of stabilization (Bayoumi and Masson 1995; Mélitz and Zumer 2002; Decressin 2002; Andersson 2008; Arachi 2010; Hepp and von Hagen 2011). I then present the cases where this manner of quantification is identical to that of Asdrubali et al. (1996) and the cases where it is not so. Given the examination of these two strands of analysis, I propose a decomposition that synthesizes the two strands in the literature and that, I believe, is more straightforward and easier to interpret. I then apply this decomposition to the NHI system in Japan.

The paper proceeds as follows. Section 2 provides an institutional background of the Japanese public health insurance in general, and the NHI system in particular. Section 3 formally elaborates on the method that quantifies the smoothing effects after reviewing the methods by the previous studies. Section 4 implements the decomposition analysis to examine the stabilization effects of fiscal transfers, and discusses the results. Finally, Section 5 presents the conclusions of the study.

2. Institutional Background

In Japan, every person receives standardized medical services at identical prices, regardless of the type of public health insurance he or she is enrolled in. There are *no* gatekeepers hindering the choice of medical services. The insured are free to choose among the many medical service providers (clinics or hospitals, private or public) regardless of the providers' locations, facility types, or other factors such as having a referral or not. The insured pay 30% of the medical fees as co-payments.[1] Payments to the providers are mainly fee-for-service.[2] The insurers reimburse 70% of the insurance-

[1] The co-payments rate for high-income persons aged 70 years and above is also 30%, but that for the other elderly is 20%.
[2] The 2006 Reform, however, introduced a package payment system for the treatment of the elderly to circumvent the possible adverse effects of this fee-for-service system.

covered service costs to the providers. Each month, bills on medical treatments and drugs are examined by the Social Insurance Medical Fee Payment Fund for the employment-based public health insurance and the National Health Insurance Federation for the NHI.[3] The range of medical services to be covered, rates of co-payment, and fee schedules of medical services are identical and standardized by law throughout the nation for all insurance programs. These standardized elements are reviewed every two years.

The public health care in Japan may be broadly categorized as employment-based public health insurance, residence-based public health insurance, and another residence-based scheme especially for the elderly. **Table 1** summarizes the schemes along with the total population covered by each system in 2009. The employment-based insurance is called the Employees' Health Insurance (EHI) in general, and is further categorized into the (i) association-managed health insurance for employers in large companies, (ii) Japan Health Association Insurance (JHAI) for employers in small and medium enterprises (SMEs), (iii) Mutual Aid Association Health Insurance for school and public sector employees, and (iv) Seamen's Health Insurance managed by the JHAI. These insurances cover employees and their dependents. On the other hand, the NHI is a residence-based scheme for those aged below 75 years and are excluded from the EHI. The NHI is further divided into two categories: the municipal NHI whose insurers are municipalities, and the NHI associations for professionals such as doctors and lawyers.

There is an additional scheme for the elderly population. The NHI previously covered the entire elderly population unless they were dependents of EHI subscribers until 1983, when a new financing scheme, the Elderly Health Care Service (EHCS), was introduced to enable municipalities to disburse medical benefits for all those aged 70 years or above (and those aged between 65 and 70 years who are bedridden or seriously disabled). The benefits are financed by transfers from central and local governments, those from the public health care insurers, and co-payments by the elderly. The elderly continued to be enrolled in their previous social insurance programs by paying their premiums until 2008, when the current Elderly Health Care Service for the Old-Old (EHCSOO) started. In 2008, the EHCSOO began separating those aged 75 years and above (i.e., the "old-old") from the public health insurance schemes. All municipalities

[3] While their local offices are supposed to inspect all bills, their capacity is limited. Intensive reviews conducted by medical experts are limited only to high-cost cases or specified suspicious facilities.

within a prefecture formed an organization that disburses the medical benefits to the old-old, financing them from premiums paid by the old-old (10%) and fiscal transfers from the public health care insurers (40%) and central and local governments (100/3% by the center; 25/3% by the prefecture; and 25/3% by the municipalities).

Table 1

The focus of the current study is the municipal NHI. Since the municipal NHI is a region-based scheme that covers those who are excluded from the EHI, the insured typically include riskier and less wealthy groups in the population, including the retired, those aged less than 75 years, self-employed non-professionals, farmers, employees of unincorporated business with less than 6 full-time employees, college students aged 20 years and above, and the unemployed. The insurers are the municipalities. Every municipality sets up and manages its own NHI insurance association. Revenues and expenditures in the NHI are accounted for in a special account apart from the municipal general account.

The benefits of a municipal NHI are financed from premiums from its subscribers and a variety of fiscal transfers. While such transfers total more than 20 items, they may be broadly categorized into those from the (i) central government (central transfers),[4] (ii) prefectural government (prefectural transfers),[5] (iii) cost-sharing scheme among all public health care insurers (net transfers from inter-institutional cost-sharing scheme),[6] (iv) cost-sharing scheme among municipalities within a prefecture (net transfers from within-prefecture cost-sharing scheme),[7] and the general account of the municipality (within-municipal transfers). The last item is further categorized into (v) transfers

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[4] The central government matches 34% of medical benefits, 25% of catastrophic medical expenses, and 33% of designated health checks and health promotion. It also provides lump-sum birth allowance. In addition, there is a Fiscal Adjustment Grant (FAG) that covers about 9% of medical benefits, and has two components: the Ordinary FAG (80% of the total FAG), which intends to equalize fiscal capacities among municipalities, and the Special FAG, which saves the remaining 20% for unexpected health care demands.

[5] Prefectural governments match 25% of catastrophic medical expenses, and 33% of health promotion expenses. They also provide the Prefectural Adjustment Grant (PAG), which is analogous to the FAG, and which constitutes 7% of relevant medical costs. Six percent of the PAG is distributed according to a fixed rule, and the remaining 1% is used for unexpected local demands.

[6] The insurers in the NHI and the EHI contribute funds to the EHCSOO based on the size of their subscribers. An equalizing scheme is also in place for the expenses of those aged between 65 and 75 years (‘the young-old’) in each public health insurance insurers. Furthermore, the Medical Benefit Grants are disbursed for the retired employees enrolled in the NHI.

[7] There are two types of cost sharing. The Catastrophic Medical Expense Grants cover expenses above JPN¥80,000 per receipt. The Fiscal Stabilization Grants also covers expenses above JPN¥300,000 per receipt.

specified by national laws[8] (statutory municipal transfers), and (vi) transfers disbursed at the municipality's discretion (discretionary municipal transfers). Note that the *net* transfers from the cost-sharing schemes of (iii) and (iv) can be negative.

Municipalities have discretions over their premium schedules. They are usually based on factors such as household income and asset, and household size and composition. Municipalities arguably set their premium schedules so that their NHI special accounts close, given the fiscal transfers as well as the other revenue items. Under the current system, medical benefits are defined by the medical needs of the insured, which clearly municipalities cannot either ration or reduce, for example, simply due to deficits in the NHI special accounts.[9] While the premium schedules are fixed within a fiscal year, they are likely to change over time in response to the fiscal status of the special accounts.

I examine how regional medical risks, reflected in changes in the NHI medical benefits, are smoothed. The key relation is between the medical benefits and the premiums within a municipality. For example, assume a change in regional medical needs. If there were no revenue sources other than the premiums, municipalities would have to change their premiums to match the change in medical benefits caused by the change in regional medical needs. Then, they must introduce multiple layers of fiscal transfers to the municipal NHI, which should help the municipalities control the changes in the premiums caused by the changes in local medical needs. I thus pay a special attention to the roles of fiscal transfers in smoothing premium payments.

For every municipality in each fiscal year, the municipal NHI accounts allow us to derive the following relation:

$$\begin{aligned}\text{Medical benefits} \equiv\ & \text{Central transfers} + \text{Provincial transfers} + \text{Net transfers from inter-} \\ & \text{institutional cost sharing} + \text{Net transfers from within-province} \\ & \text{cost sharing} + \text{Statutory municipal transfers} + \text{Discretionary} \\ & \text{municipal transfers} + \text{Other net revenues} + \text{Inter-temporal} \\ & \text{adjustments} + \text{Premiums}. \qquad \text{(1a)}\end{aligned}$$

[8] Transfers from the general budgets prescribed by national laws have several categories. The first is for premiums abatement measures for low-income subscribers, which is further categorized into three subcategories. The second is for medical needs that are not controlled by the insurers. There are also subsidies for lump-sum birth allowance and management costs.

[9] A health promotion measure may influence medical needs in the long run, but not in the short term.

where “inter-temporal adjustments” consist of “provisions from reserves,” “debt issues,” and “surpluses,” net of “addition to reserves” and “debt-service payments.” In addition, “other net revenues” is the remaining miscellaneous revenues minus the remaining miscellaneous expenditures. The following section explains how I measure the smoothing effects of the items described above, improving on the previous studies on the stabilization effects among regions.

3. Measuring Stabilization Effects

Let x_{it} be the “risk” variable of region i in year t. In the current case, this is the per capita NHI medical benefits to which shocks occur. Assume that the risk variable x_{it} can be decomposed into J elements $\{y_{j,it}\}$ as

$$x_{it} = \sum_{j=1}^{J-1} y_{j,it} + y_{J,it} \tag{1b}$$

where $y_{J,it}$ is the “target” variable, an outcome variable brought about as a smoothed-risk variable. The identity (1b) allows us to decompose the variance of x_{it} into indicators for the degree of contributions of $y_{j,it}$ for $j = 1, \ldots, J–1$ in buffering $y_{J,it}$ from shocks in x_{it}. Note that (1b) is comparable to (1a) where x_{it} and $y_{J,it}$ are respectively NHI benefits and premiums, both in per subscriber terms.

2.1. Standard decomposition using differenced-log variables

I start the discussion with the method developed by Asdrubali et al. (1996), which has been utilized by, among others, Doi (2000), Buettner (2002), Andersson (2004), Jüßen (2006), Ramos and Coimbra (2009), Furceri (2010) and Balli et al. (2011). Using Asdrubali et al.’s (1996) method, these studies typically decomposed the variance of gross products (risk variable) into its covariance with elements of the system of economic accounts, and measured the contributions of such elements in preventing personal disposable income or consumption (target variable) from shocks in the gross products. The method takes advantage of the following identity:

$$x_{it} \equiv \frac{x_{it}}{x_{it} - y_{1,it}} \cdot \frac{x_{it} - y_{1,it}}{x_{it} - \sum_{j=1}^{2} y_{j,it}} \cdot \frac{x_{it} - \sum_{j=1}^{2} y_{j,it}}{x_{it} - \sum_{j=1}^{3} y_{j,it}} \cdot \ldots \cdot \frac{x_{it} - \sum_{j=1}^{J-2} y_{j,it}}{x_{it} - \sum_{j=1}^{J-1} y_{j,it}} \cdot y_{J,it} \tag{2}$$

where $y_{J,it} = x_{it} - \sum_{j=1}^{J-1} y_{j,it}$ by (1b). Note that $y_{j,it}$ can be negative or positive for $j \neq J$ but $x_{it} - \sum y_{j,it}$ and $y_{J,it}$ are always positive. Taking the log of (2) and differencing the resulting terms yield

$$\begin{aligned}\Delta \ln x_{it} &= \Delta \ln\left[x_{it} / \left(x_{it} - y_{1,it}\right)\right] \\ &+ \sum_{k=1}^{J-2} \Delta \ln\left[\left(x_{it} - \sum\nolimits_{j=1}^{k} y_{j,it}\right) / \left(x_{it} - \sum\nolimits_{j=1}^{k+1} y_{j,it}\right)\right] + \Delta \ln y_{J,it}.\end{aligned} \tag{3}$$

Then, (i) subtracting the expectations of each term in (3), (ii) multiplying both sides by $\{\Delta\ln x_{it} - \mathrm{E}[\Delta\ln x_{it}]\}$, (iii) taking other expectations of the resultant products, and (iv) dividing both sides by $\mathrm{var}[\Delta\ln x_{it}]$ result in

$$\begin{aligned}1 &= \frac{\mathrm{cov}\left\{\Delta \ln x_{it}, \Delta \ln\left[x_{it} / \left(x_{it} - y_{1,it}\right)\right]\right\}}{\mathrm{var}\left(\Delta \ln x_{it}\right)} \\ &+ \sum_{k=1}^{J-2} \frac{\mathrm{cov}\left\{\Delta \ln x_{it}, \Delta \ln\left[\left(x_{it} - \sum\nolimits_{j=1}^{k} y_{j,it}\right) / \left(x_{it} - \sum\nolimits_{j=1}^{k+1} y_{j,it}\right)\right]\right\}}{\mathrm{var}\left(\Delta \ln x_{it}\right)} \\ &+ \frac{\mathrm{cov}\left(\Delta \ln x_{it}, \Delta \ln y_{J,it}\right)}{\mathrm{var}\left(\Delta \ln x_{it}\right)}.\end{aligned} \tag{4}$$

Each term on the right-hand side of (4) is an ordinary least squares (OLS) estimator for a coefficient on $\Delta\ln x_{it}$ in a linear model that regresses $\Delta\ln[(x_{it} - \sum^{k}_{j=1} y_{j,it})/(x_{it} - \sum^{k+1}_{j=1} y_{j,it})]$ on a constant and $\Delta\ln x_{it}$. In particular, the last term is a coefficient on $\Delta\ln x_{it}$ from the OLS regressing $\Delta\ln y_{J,it}$ on a constant and $\Delta\ln x_{J,it}$. This term must be zero if the target variable does not correlate with the risk variable, showing that the variations in the risk variable are perfectly absorbed by the terms other than $\Delta\ln x_{J,it}$. It is then natural to define the index of smoothing (or stabilization) as

$$\begin{aligned}1 - \gamma &= \frac{\mathrm{cov}\left\{\Delta \ln x_{it}, \Delta \ln\left[x_{it} / \left(x_{it} - y_{1,it}\right)\right]\right\}}{\mathrm{var}\left(\Delta \ln x_{it}\right)} \\ &+ \sum_{k=1}^{J-2} \frac{\mathrm{cov}\left\{\Delta \ln x_{it}, \Delta \ln\left[\left(x_{it} - \sum\nolimits_{j=1}^{k} y_{j,it}\right) / \left(x_{it} - \sum\nolimits_{j=1}^{k+1} y_{j,it}\right)\right]\right\}}{\mathrm{var}\left(\Delta \ln x_{it}\right)}.\end{aligned} \tag{5}$$

where $\gamma \equiv \mathrm{cov}(\Delta\ln x_{it}, \Delta\ln y_{it})/\mathrm{var}(\Delta\ln x_{it})$. This shows that the smoothing effect $1 - \gamma$ is decomposed into a series of covariance-variance ratios that appear in the right-hand side of (5), each of which shows its relative contribution in buffering the target variable from the risk variable. The differenced log ratio is the difference between the growth rates of

$(x_{it} - \sum^{k}_{j=1} y_{j,it})$ and $(x_{it} - \sum^{k+1}_{j=1} y_{j,it})$, which is interpreted in the literature as the growth rates attributable to $y_{k+1,it}$.

In the previous studies, the order of the components of x_j is obvious as they follow the system of economic accounts. For example, consider a case of three components of personal income x, y_1, y_2, and y_3, where $x - y_1$ is disposal income and $y_3 = x - y_1 - y_2$ is consumption. This relation then makes y_1 net taxes and y_2 net savings, and defines their ordering. Then, (5) implies the following regressions:

$$\begin{aligned} \Delta \ln(x) - \Delta \ln(x - y_1) &= \alpha_1 + \beta_1 \cdot \Delta \ln x + u_1 \\ \Delta \ln(x - y_1) - \Delta \ln(x - y_1 - y_2) &= \alpha_2 + \beta_2 \cdot \Delta \ln x + u_2 \text{ and} \\ \Delta \ln y_3 &= \alpha_3 + \beta_3 \cdot \Delta \ln x + u_3 \end{aligned} \tag{6}$$

where the *it* subscripts are dropped for expositional convenience. The estimates for β_1 and β_2 are interpreted to capture the buffering effects of y_1 and y_2, respectively, and that for β_3 is used to construct the risk-sharing effect as $1 - \beta_3$.

This decomposition, however, has a problem when the elements have no natural order among the items. For example, I could use the following regressions to estimate ϕ_1 and ϕ_2, which respectively capture the buffering effects of y_2 and y_1:

$$\begin{aligned} \Delta \ln(x) - \Delta \ln(x - y_2) &= \delta_1 + \phi_1 \cdot \Delta \ln x + u_1 \\ \Delta \ln(x - y_2) - \Delta \ln(x - y_2 - y_1) &= \delta_2 + \phi_2 \cdot \Delta \ln x + u_2 \text{ and} \\ \Delta \ln y_3 &= \delta_3 + \phi_3 \cdot \Delta \ln x + u_3. \end{aligned} \tag{7}$$

Except the last regression model, the OLS estimates for ϕ_1 and ϕ_2 in (7) are generally different from those for β_1 and β_2 in (6). This then raises a concern that the standard decomposition method by Asdrubali et al. (1996) will not work for the current case where the order of the components of a risk variable is not "obvious" or "natural."

3.2. Decomposition in differenced level

Since the items in the NHI account do not have a predetermined order in the same way that the system of national accounts imply as above, I have to forego the popular method and utilize an alternative decomposition that is independent of the order of the items in the NHI accounts. In fact, the decomposition I propose is found to be more straightforward and easier to interpret.

The procedure is analogous to the standard decomposition, but dispenses with (2) and only utilizes (1b) by differencing it as

$$\Delta x_{it} \equiv \sum_{j=1}^{J-1} \Delta y_{j.it} + \Delta y_{J,it} \tag{8}$$

which yields the deviations from expectations

$$\Delta x_{it} - E(\Delta x_{it}) \equiv \sum_{j=1}^{J-1} [\Delta y_{j.it} - E(\Delta y_{j.it})] + [\Delta y_{J,it} - E(\Delta y_{J,it})].$$

Then, multiplying this expression with [Δx_{it} − E(Δx_{it})], and taking the expectation of the resulting products and dividing the expectations by var(Δx_{it}), I obtain

$$1 = \sum_{j=1}^{J-1} \frac{\operatorname{cov}\left(\Delta x_{it}, \Delta y_{j,it}\right)}{\operatorname{var}\left(\Delta x_{it}\right)} + \frac{\operatorname{cov}\left(\Delta x_{it}, \Delta y_{J,it}\right)}{\operatorname{var}\left(\Delta x_{it}\right)}. \tag{9}$$

Analogous to (5), each term on the right-hand side of (9) is obtained as an estimate for a coefficient on Δx_{it} from an OLS regression of $\Delta y_{j,it}$ on a constant and Δx_{it}. Again, if the target variable does not correlate with the risk variable where shocks occur, the last term will be zero and the variations in the risk variable are perfectly absorbed by the terms other than $y_{J,it}$ in (8). Then, the index of stabilization here is defined as

$$1 - \beta_J = \sum_{j=1}^{J-1} \frac{\operatorname{cov}\left(\Delta x_{it}, \Delta y_{j,it}\right)}{\operatorname{var}\left(\Delta x_{it}\right)} = \sum_{j=1}^{J-1} \beta_j . \tag{10}$$

where $\beta_j \equiv$ cov(Δx_{it}, $\Delta y_{j,it}$)/var(Δx_{it}). Now, the interpretation is simpler and more intuitive. The stabilization effect $1 - \beta_J$ is decomposed into β_js for $j = 1,\ldots, J - 1$, which are the covariance between changes in the risk variable and those in each of the components that exclude the target variable $y_{J,it}$. In addition, the β_j coefficients are independent of the order of $y_{j,it}$ for $j = 1, \ldots, J–1$.

3.3. Comparison and synthesis with the "ad-hoc" method

Keeping (5) and (10) in mind, I examine another strand in the literature on "risk sharing" among regions. In a rather ad-hoc manner, this literature typically regressed *some* form of the target variable $y_{J,it}$ (i.e., regional gross output) on a constant and some form of the risk variable x_{it} (i.e., regional consumption) to obtain the coefficient estimate θ on the latter variable (Bayoumi and Masson 1995, Mélitz and Zumer 2002, Decressin 2002, Andersson 2008, Arachi et al. 2010; Hepp and von Hagen 2011). It then argued that $1 - \theta$ indicates the degree of risk sharing.

This manner of estimation may or may not be exactly based on a foundation like (5) or (10). In particular, the specific form of risk and target variables varies among the studies. For example, Andersson (2008) uses a differenced *log* of variables but normalizes the pre-logged variables by their national averages to net out an aggregate shock in a given year. In this case, the OLS estimate for θ is identical to that for γ in (5) *if the decomposition uses cross-section data*. Let us define $a_t \equiv 1/N^{-1}\sum_i x_{it}$ and $b_t \equiv 1/N^{-1}\sum_i y_{J,it}$, where $N^{-1}\sum_i y_{J,it}$ and $N^{-1}\sum_i x_{it}$ are national averages with N being the number of regions. It then follows that

$$\frac{\operatorname{cov}\left(\Delta\ln a_t x_{it}, \Delta\ln b_t y_{J,it}\right)}{\operatorname{var}\left(\Delta\ln a_t x_{it}\right)} = \frac{\operatorname{cov}\left(\Delta\ln x_{it} + \Delta\ln a_t, \Delta\ln y_{J,it} + \Delta\ln b_t\right)}{\operatorname{var}\left(\Delta\ln x_{it} + \Delta\ln a_t\right)}$$

$$= \frac{\operatorname{cov}\left(\Delta\ln x_{it}, \Delta\ln y_{J,it}\right)}{\operatorname{var}\left(\Delta\ln x_{it}\right)} = \gamma.$$

because $\Delta\ln a_t$ and $\Delta\ln b_t$ take on common fixed values across the regions in a given single period t. In other words, normalizing the variables with their national averages does not change the covariance with a cross section of data. By referring to (4) and (5), we then see that the ad-hoc estimate for θ *with a cross section of differenced logged variables* is identical to the OLS estimate for γ by Asdrubali et al. (1996).

On the other hand, however, other studies (Bayoumi and Masson 1995, Mélitz and Zumer 2002, Decressin 2002, and Arachi 2010) use differenced *levels* of variables normalized by their averages, that is, $\Delta(y_{J,it}/N^{-1}\sum_i y_{J,it})$ and $\Delta(x_{j,it}/N^{-1}\sum_i x_{j,it})$, to estimate θ. In this case, even with a cross section of data, the OLS estimate for θ is *not* identical to that for β_J in (10), the decomposition the current study proposes. This is because

$$\frac{\operatorname{cov}\left(\Delta a_t x_{it}, \Delta b_t y_{J,it}\right)}{\operatorname{var}\left(\Delta a_t x_{it}\right)} = \frac{\operatorname{cov}\left(a_t x_{it} - a_{t-1}x_{it-1}, b_t y_{J,it} - b_{t-1}y_{J,it-1}\right)}{\operatorname{var}\left(a_t x_{it} - a_{t-1}x_{it-1}\right)} \neq \frac{\operatorname{cov}\left(\Delta x_{it}, \Delta y_{J,it}\right)}{\operatorname{var}\left(\Delta x_{it}\right)} = \beta_J$$

which is easily intuited since (1b) does not hold if its elements are normalized by their *respective* national averages. I argue, however, that we do not need to normalize the variables by their national averages even if we want to net out an *aggregate* shock in a given year. This is because the constant term with a cross section of data can control an aggregate shock in the same way that time dummies take care of such shocks in a panel-data analysis. In fact, if the level variables are *not* normalized by their respective national averages, we find a theoretical foundation in (10) for the seemingly "ad-hoc" estimate for the stabilization effect, since $\beta_J \equiv \text{cov}(\Delta x_{it}, \Delta y_{J,it})/\text{var}(\Delta x_{it})$.

4. Empirical Implementation

In the following empirical implementation, I thus utilize the decomposition (10) with a cross section of differenced data for *each* of fiscal years from FY 2003–2009. For the reason I argued in Section 3.1, I do not use the decomposition with differenced logged data as the previous studies had done so far. In addition, for the reason I explained in Section 3.2, I do not normalize the variables in level by their respective national averages either.

4.1. Sample and Data

I obtained all the data from the municipal special accounts for the NHI for FY 2002–2009. However, the data that are aggregated at the prefectural level are only publicly available, except those for FY 2008 and FY 2009. I then took advantage of the Access to Government Information Act to obtain the municipal data for FY 2002–2007 directly from the Ministry of Health, Labour and Welfare (MHLW). I then used the municipal data to construct the variables that appear in (1a), namely (a) medical benefits, (b) central transfers, (c) provincial transfers, (d) net transfers from within-province cost sharing, (e) net transfers from inter-institutional cost sharing, (f) statutory transfers from municipal general accounts (hereafter "statutory municipal transfers"), (f) discretionary transfers from municipal general accounts (hereafter "discretionary municipal transfers"), (g) net other revenues, (h) inter-temporal adjustments, and (i) premiums. All variables are expressed in per-subscriber terms. In addition, the variables are not deflated by a price index because the estimation utilizes a cross section of data.

Table 2 shows the shares of elements (b)–(i) among the medical benefits from FY 2002–2009. The shares of premiums, central transfers, and net transfers from inter-institutional cost-sharing schemes occupy larger portions, although the sign of the last share is negative for FY 2002–2007. At the same time, their values have changed over time. As time passed, while the shares of premiums and central transfers have decreased, the share of net transfers from inter-institutional cost-sharing schemes has increased. In addition, the share of provincial transfers has also increased from 0.4% in FY 2002 to over 6% after FY 2007. These are apparently due to institutional changes that happened almost every year during the period as listed in Table 3. Note that in Table 2, the share of net transfers from within-prefecture cost sharing tends to be small since these values are obtained as net values aggregated at the national level.

Tables 2 and 3

For the estimation, I use a cross section of differenced values of the medical benefits as x_{it} and each of the nine variables listed in Table 2 as $y_{j,it}$. The sample statistics for the nine annually differenced variables are listed in Table 4, along with those for the all-period observations.

In contrast to the current study, the previous empirical studies utilized a panel data of regions, assuming constant coefficients or covariance-variance ratios over the period.[10] This may be because, with the exception of Andersson (2004) and Jüßen (2006), who respectively used samples of 279 municipalities in Sweden and 439 counties in Germany, they examined cases of the highest layer of sub-national government whose numbers tend to be too small for conducting a reasonable cross-section analysis.[11] On the other hand, the sample sizes of differenced cross-sectional units in this study are over 1,700, which should be large enough for performing a reasonable cross-sectional analysis.

In addition, it may not be appropriate to assume that the covariance-variance ratios are constant over the periods for the current study, since such ratios are likely to have changed almost every year for the following reasons. First, due to the mergers in the 2000s, which peaked in 2004 and 2005, the number of municipalities has decreased from about 3,200 to 1,700. These changes in the number of municipalities plausibly affect the stabilization effects of the channels of the NHI financing. In addition, such mergers are interpreted as elements of shocks to local health demand in municipalities

[10] The only exception is Asdrubali et al. (1996), who, in addition to a constant-coefficient panel analysis, also performed a cross-sectional analysis with 48 contiguous US states and examined how channels of risk sharing change over time, as I do in this study.

[11] For example, there are 10 provinces in Canada (Bayoumi and Masson 1995, Mélitz and Zumer 2002, Balli et al. 2011), 21 regions in France (Mélitz and Zumer 2002), 16 Länders in Germany (Buettner 2002, Hepp and von Hagen 2011), 20 regions in Italy (Decressin 2002, Arachi et al. 2010), 47 prefectures in Japan (Doi 2000), 30 regions in Portugal (Ramos and Coimbra 2009), 21 regions in Sweden (Andersson 2008), 21 regions in the UK (Mélitz and Zumer 2002), and 50 states in the US (Asdrubali et al. 1996, Bayoumi and Masson 1995, Mélitz and Zumer 2002). In practice, most of the studies utilized even smaller number of regions than those listed here due to data limitations.

that have annexed other municipalities, since their medical benefits per subscriber must have changed due to such mergers (annexations). However, I had no choice but to exclude municipalities that merged to form *new* municipalities or were annexed by other municipalities in period t from the sample for that year, since I use a cross section of differenced variables over two adjacent fiscal years, t and $t - 1$.

Second, there has been a series of institutional changes in the NHI in the 2000s as shown in Table 3, which mainly, but not exclusively, includes changes in the transfer systems. These changes affected the *difference* data for every period in the 2000s, since they affected the variables *in level* for in period t or period $t - 1$, or in both periods t and $t - 1$ that constitute the corresponding variables *in difference*.

Table 4

4.2. Estimation and Econometric Issues

The regression model that I estimate is therefore

$$\Delta y_{j.it} = \alpha_j + \beta_j \cdot \Delta x_{it} + u_{it} \tag{11}$$

for a given single fiscal year t that runs from 2003 to 2009. Subscript i indexes municipalities (insurers) and j indexes each of the eight components (all in per-subscriber values) of the right-hand side of identity (1a). What we are interested in are the estimates for β_j and their change over time.

There are two views on how to conceptualize (11). The first view regards (11) as the data generating process (GDP) for Δy_j, while the second view considers (11) a convenient *artificial* regression to produce the covariance-variance ratios defined in (9) or (10). In this current study, I am more inclined to take the second view, while I argue that the analogous studies on the nexus between gross product and disposable income (or consumption) can be considered as subscribing to the first view.

If the first view is taken, it is understandable that the literature concerns the endogeneity of the regressors (e.g., Asdrubali et al. 1996, Andersson 2004), since they typically regressed a target variable that includes taxes or transfers on a constant and a risk variable (gross product or personal income). It is indeed plausible that taxes or transfers are determined by gross product or personal income, causing a reverse causation and endogeneity. However, even if I took the first view, the current case is somewhat different from the previous studies. I regress each of the eight fiscal variables on the medical benefits. Since the medical benefits depend on the health status of the insured, it would be hard to imagine that causation runs from the fiscal variables to the health status or medical benefits. Still, it might be possible that, for example, transfers for local health promotion expenditures improve health status, thereby decreasing medical benefits. However, even if they do, it should take time for the effects to be realized. Therefore, unless the error term in regression model (11) has a long memory, the endogeneity would not be much of a concern in terms of reverse causation.

However, there may be another type of endogeneity caused by unobserved heterogeneity. If there are unobserved elements that affect the dependent variable and are correlated with the observed regressor, then the error term may be correlated with the observed regressor. Note nonetheless that (11) is a *static* first-differenced model where the unobserved heterogeneity has already been differenced if it affects only the dependent variable in level and is, as conventionally assumed, constant over the period. However, differencing a regression model results in serial correlation in the differenced error term $u = e_t - e_{t-1}$, unless e_t, the error term *before* differencing, follows the unit-root process $e_t = e_{t-1} + u_t$, and u_t is serially independent. Therefore, again, if we believe that the regressor is independent of the short-lagged error term u and that the error term in fact has a short memory, the endogeneity due to the first-differencing would not be much of a concern.

This issue, if we continue to hold the first view of (11), relates to another issue of the non-spherical error term. For example, Asdrubali et al. (1996) discussed the possibility that variances of the error term change over time across the cross-sectional units. With a panel data, I could have allowed for such an issue by the method of clustering. However, I use a cross section of data. Although the feasible generalized least squares (FGLS) may still be applicable to the current analysis, the functional form of the conditional variance of the error term is rarely known in practice, as argued by Stock and Watson (2010). If the functional form is not correct, the FGLS standard errors are invalid in the sense that they lead to incorrect statistical inferences. Therefore, when I take the first view of (11) at all, I use OLS to estimate the βs along with the heteroskedasticity-consistent standard errors.

The second view regards (11) as an *artificial* regression to produce the covariance-variance ratios in (9) or (10), in which we aim to obtain the descriptive statistics for cov(Δx, Δy_j)/var(Δx). What we want to obtain is cov(Δx, Δy_j)/var(Δx), not the other quantities, since they are defined as such in the analytical formulation of (9) or (10). Regression model (11) is thus a convenient tool for obtaining $\beta_j \equiv$ cov(Δx, Δy_j)/var(Δx), in which I emphasize identity "≡." With this definition, only OLS can be applied to (11) to obtain the relevant quantities. Furthermore, this second view does not allow other methods such as FGLS or instrumental variable (IV) estimators, since they yield quantities for β_j that are constructed differently from cov(Δx, Δy_j)/var(Δx). This study is more inclined to take this second view, since the derivation of the quantities in question is based on (9) or (10). In this case, the discussion of endogeneity or non-spherical errors would be irrelevant. However, in what follows, I, in effect, also take the first view when I discuss the standard errors and confidence intervals that are obtained from the OLS on (11).

4.3. Results and Interpretation

The estimation results are shown in Table 5 and Figure 1. Figure 2 contains nine panels that show the changes in the coefficients over time for each of the nine elements, with a 95% confidence interval. The results are summarized in the following.

Table 5 and Figures 1 and 2

First, the very small percentages of the premiums throughout the periods indicate that the stabilizing effect (10) is very large, albeit with some variations over the periods. Notice that they are negative except in FY 2003 and FY 2008. While the negative value implies over-smoothing in that a shock to the medical benefits reduces the change in the premiums, in the first four and the last periods, both the negative and positive coefficients are statistically insignificant, implying that the effect is nil. In the other two periods (FY 2007 and FY 2008), the values are statistically significant. In FY 2007, the value (–2.0) implies over-smoothing, albeit it is small. In FY 2008, the value (3.8%) is the largest throughout the period, which may be due to a major institutional change, as I explain below. I therefore argue that the degree of risk sharing or the buffering effects on the premiums from changes in medical demands has been almost perfect during the period.

Second, this high degree of risk-sharing effects can be attributed to the combined effects of the transfer payments. This is shown by the relatively low values for the items other than the transfers, namely, the inter-temporal adjustments and the other net revenues. However, the inter-temporal adjustments exhibit as high a coefficient value as 20% in the first two and the last periods. The contributions of the other net revenues are smaller. The latter coefficients are not statistically significant for the first four periods, implying that their smoothing effects are nil. For the last three periods, the item's contribution is at most 9.2% with a negative value of −2.7% in FY 2008.

Third, the effects among the transfers are different. Among the six categories, the effects of the central transfers, net transfers from inter-institutional sharing schemes, and those from within-prefectural cost-sharing schemes are relatively large and statistically significant in all the years. The remaining three types of transfers have minor effects. The effects of provincial transfers are statistically insignificant for the first five periods. Although their values are still small, the two types of municipal transfers have larger effects than the provincial transfers. In particular, the discretionary municipal transfers de-smoothed the changes in the medical benefits in the last two periods, with a relatively large negative value (−12.0) in FY 2009. This is an interesting result since it is often argued that municipalities use their discretional transfers as an ex-post instrument to cover the deficits in their NHI special accounts, and therefore they are a source of a lack of fiscal discipline in municipal fiscal management. However, the statistically negative coefficients in FY 2008–2009 indicate otherwise: if the claim were the case, the coefficient should have yielded statistically significant positive values.

Fourth, the varying values of the coefficients over the years corroborate the fact that the institutional aspects of the NHI system have changed almost every year. Such changes are salient for almost every item in FY 2008. In that year, those aged 75 years and above who had been enrolled in the NHI became covered by the HCSOO. In addition, public health care insurers including municipalities started to contribute funds to finance the medical benefits paid by the HCSOO. Furthermore, a new equalizing scheme among the public health insurers began for the medical benefits of those aged between 65 and 75 years who are mainly enrolled in the NHI. It is thus plausible that this set of major institutional changes caused the significant fluctuations in the coefficient values in FY 2008 in Figure 1.

Last, the trends in the stabilizing effects become more articulated by dropping the values for FY 2008. Since the coefficient values are obtained from the differenced data in level, the institutional change in FY 2008 is likely to affect the values only in that year, which may then obscure the long-term trends of the buffering effects. To uncover the trends, I removed the values for FY 2008 from Figure 1, as shown in Figure 3. Indeed, Figure 3 more saliently shows the changing effects of the transfers throughout the 2000s. The most noticeable effect is the increasing role of the net transfers from the within-prefectural cost-sharing schemes. Likewise, the role of the central transfers seems to be on an upward trend, although it dropped in FY 2007. In contrast, the net transfers from the inter-institutional cost-sharing schemes are on a declining trend, dropping significantly in FY 2009. Both the municipal transfers as well as the prefectural transfers are on a declining trend also.

Figures 3

5. Concluding Remarks

In this study, I measured the degree of stabilization in the NHI in Japan. In the NHI, there are currently more than 1,700 municipalities as insurers and several institutional routes to buffer asymmetric shocks in regional health demands among municipalities. In the process of conducting the empirical analysis, I elaborated on the methods utilized in the literature on regional risk sharing (i.e., stabilization or smoothing) to decipher the relation between the two strands in the risk-sharing literature. I then developed my own decomposition method, which I employed in this study. Then, using the accounting data for the municipal NHI in the 2000s, I quantified the effects of the institutional channels, which have smoothed the changes in the NHI premiums, and discussed the behaviors of each channel in the 2000s.

I found that the NHI premium collections are almost perfectly smoothed. In particular, the system of transfers in the NHI plays a major role in stabilizing the premium collections. In contrast, the inter-temporal adjustments played a rather minor role. As mentioned in the introduction, Japan is the only country that maintains a non-competitive multiple payer system based on the traditional German system. The Japanese NHI is comparable to the regional sickness funds that once existed in

Germany and the health insurance societies for the self-employed in Korea. This study implies that despite the fact that these two countries have now departed from the traditional model, Japan still follows the traditional system mainly because of the strong role of fiscal transfers. Although comparable historical studies have to be done, I suspect that the stabilization function in the NHI would be the strongest among the comparable public health insurance schemes all over the world.

Even if such fiscal transfers are effective in smoothing *changes* in local burdens, they may not necessarily be so in equalizing the *level* of the burdens. This study only examined the stabilizing effect and left out the issue of redistribution, which most of the previous studies have considered. This omission in this study is because most of the transfer payments in the NHI system are financed *outside* of the system. As such, regional redistribution is incomplete *within* the NHI system; regional redistribution in health care concerns all the insured who are enrolled not only in the NHI but also in the other public health care programs. In fact, there is much to improve in terms of regional redistribution in the Japanese health care finance (Hayashi 2010). For example, the annual municipal premium collection per subscriber ranged from JPN¥34,724 to JPN¥142,260 in FY 2009. This also implies that the horizontal equity is compromised since households enrolled in the NHI with comparable characteristics (e.g., health risk or income) face different premium levels depending on the municipality of their residence. For example, Kitaura (2007) showed that a couple with an annual income of approximately JPN¥2.3 million, depending on their place of living, faced a variety of annual premiums ranging from below JPN¥ 60,000 to above JPN¥400,000 in FY 2003. This may imply weak equalizing effects of the transfer systems in the NHI system. Clearly, however, measuring the weakness of the equalizing effect requires another empirical analysis, which can be a future topic for research.

References

Arachi, G., Ferrario, C., Zanardi, A., 2010. Regional distribution and risk sharing in Italy: The role of different tiers of government. Regional Studies 44(1), 55-62.

Andersson, L., 2004. Regional risk-sharing provided by the fiscal system: Empirical evidence from Sweden. Regional Studies 38(3), 269-280.

Andersson, L., 2008. Fiscal flows and financial markets: To what extent do they provide risk sharing with Sweden? Regional Studies 42(7), 1003-1011.

Asdrubali, P., Sorensen, B.E., Yosha, O., 1996. Channels of interstate risk sharing: United States 1963-1990. Quarterly Journal of Economics 111(4), 1081-1110.

Balli. F., Basher, S.A., Louis, R.J., 2011. Channels of risk-sharing among Canadian Province: 1961-2006. Empirical Economics (Online First) DOI 10.1007/s00181-011-0488-6.

Bärninghausen, T., Sauerborn, R., 2002. One hundred and eighteen years of the German health insurance system: Are there any lessons for middle- and low-income countries? Social Science and Medicine 54(10), 1559-1587.

Bayoumi, T., Masson, P.R., 1995. Fiscal flows in the United States and Canada: Lessons for monetary union in Europe. European Economic Review 39(2), 253-74.

Boadway, R., Bruce, N., 1980. *Welfare Economics*. Oxford, Basil Blackwell.

Boadway, R., Hayashi, M., 2004. An evaluation of the stabilization properties of equalization in Canada. Canadian Public Policy 30(1), 91-109.

Buettner, T., 2002. Fiscal federalism and interstate risk sharing: Empirical evidence from Germany. Economics Letters 74(2), 195–202.

Crucini, M.J., 1999. On international and national dimensions of risk sharing. Review of Economics and Statistics 81(1), 73-84.

Decressin, J., 2002. Regional income redistribution and risk sharing: How does Italy compare in Europe? Journal of Public Economics 86(2), 287-306.

Doi, T., 2000. *Chiho Zaisei no Seiji-Keizai Gaku* [The Political Economy of Local Public Finance] Tokyo, Toyo Keizai Shinposha. (in Japanese).

Ferrario, C., Zanardi, A., 2011. Fiscal decentralization in the Italian NHS: What happens to interregional redistribution? Health Policy 100(1):71-80.

Furceri, D., 2010. Stabilization effects of social spending: Empirical evidence from a panel of OECD countries. North American Journal of Economics and Finance 21(1), 34-48.

Hayashi, M., 2010. "Social Protection in Japan: Current State and Challenges," In: Asher, M.G. Oum, S, Parulian, F. (Eds.) *Social Protection in East Asia: Current State and Challenges* (Economic Research Institute for ASEAN and East Asia, Jakarta) 19-54.

Hepp, R., Von Hagen, J., 2011. Fiscal federalism in Germany: Stabilization and redistribution before and after unification. Publius 42(2), 234-259.

Hussey, P., Anderson, G.F., 2003. A comparison of single- and multi-payer health insurance systems and options for reform. Health Policy 66(3), 215-228.

Jüßen, F., 2006. Interregional risk sharing and fiscal redistribution in unified Germany. Papers in Regional Science 85(2), 235-255.

Kitaura, Y., 2007. Kenko hokenryo (zei) no suiheiteki fubyodosei [Horizontal equity in the National Health Insurance Premiums]. KISER Discussion Paper Series No.8, 2007/9.

Kwon, S., 2003. Healthcare financing reform and the new single payer system in the Republic of Korea: Social solidarity or efficiency? International Social Security Review 56, 75-94.

Light, D., 2000. Sociological perspectives on competition in health care. Journal of Health Politics, Policy and Law 25(5), 696-674.

Mélitz, J., Zumer, F., 2002. Regional redistribution and stabilization by the center in Canada, France, the UK and the US: A reassessment and new tests. Journal of Public Economics 86(2), 263-86.

Ramos, P. N., Coimbra, C., 2009. The income inter-regional redistribution and the income spatial stabilization effects: An application to Portugal. Regional Science Inquiry Journal 1(1), 35-44.

Figure 1. Trends of decomposed elements

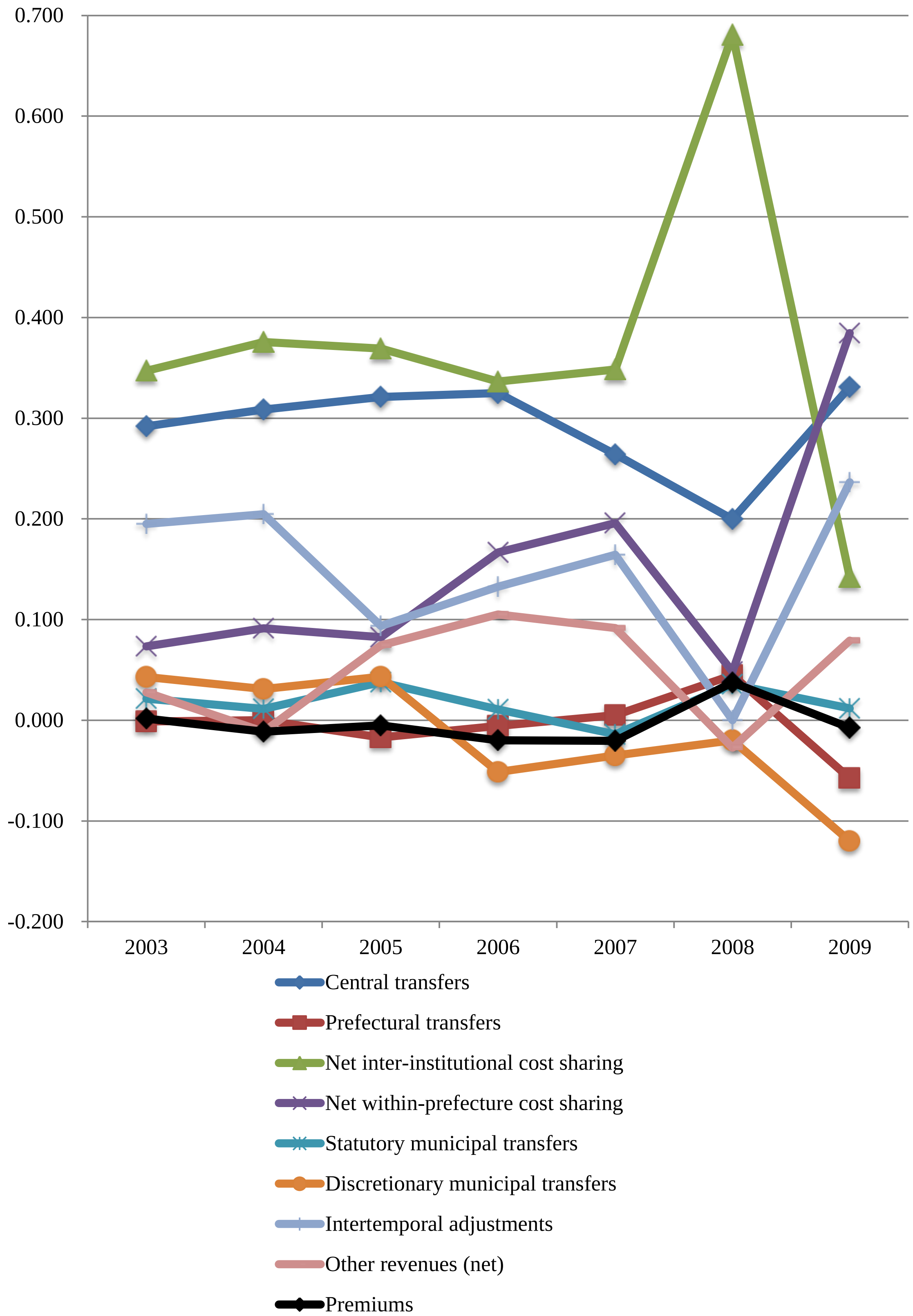

Figure 2. Trends of decomposed elements with 95% confidence intervals

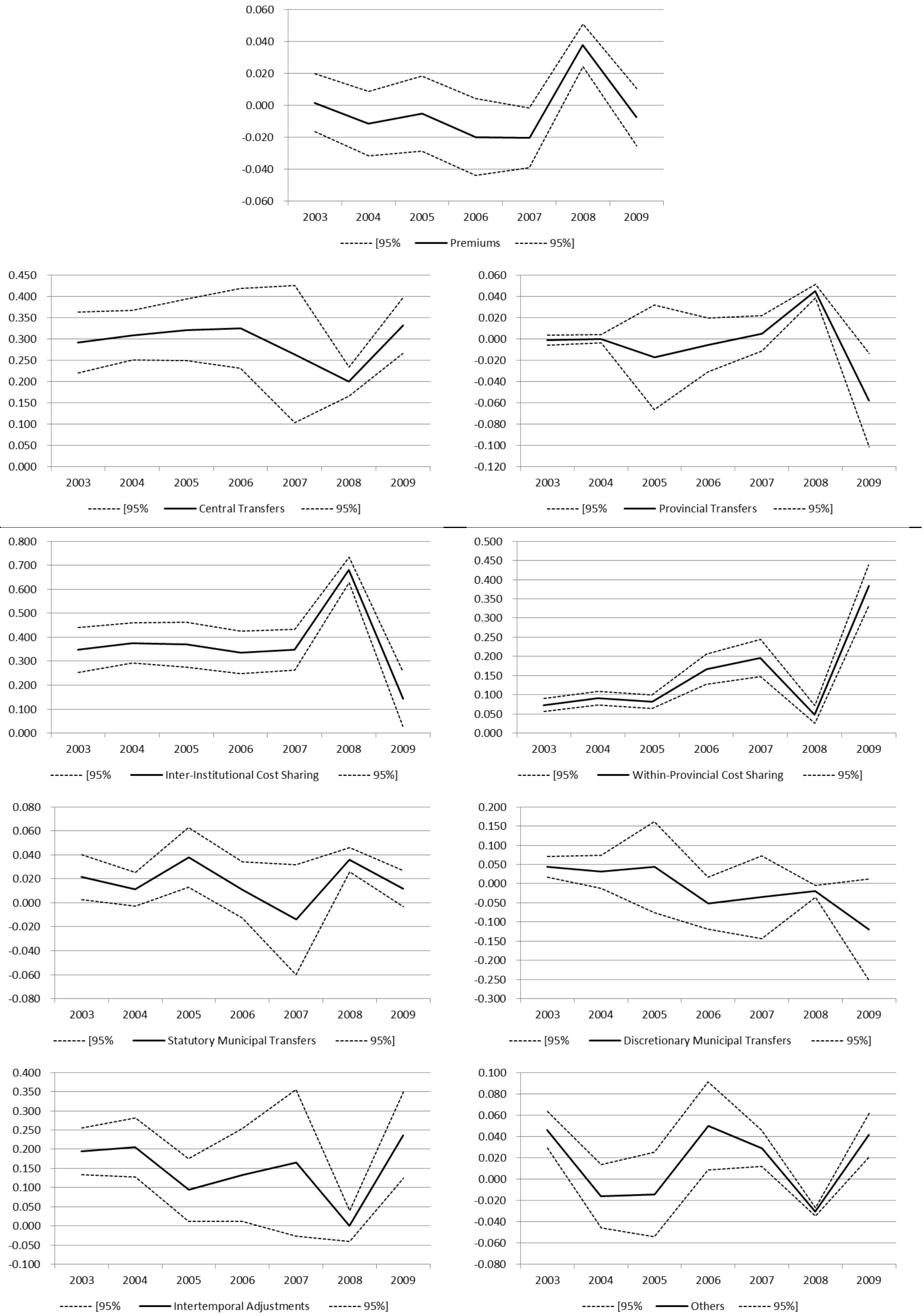

Figure 3. Trends of decomposed elements without FY 2008 values

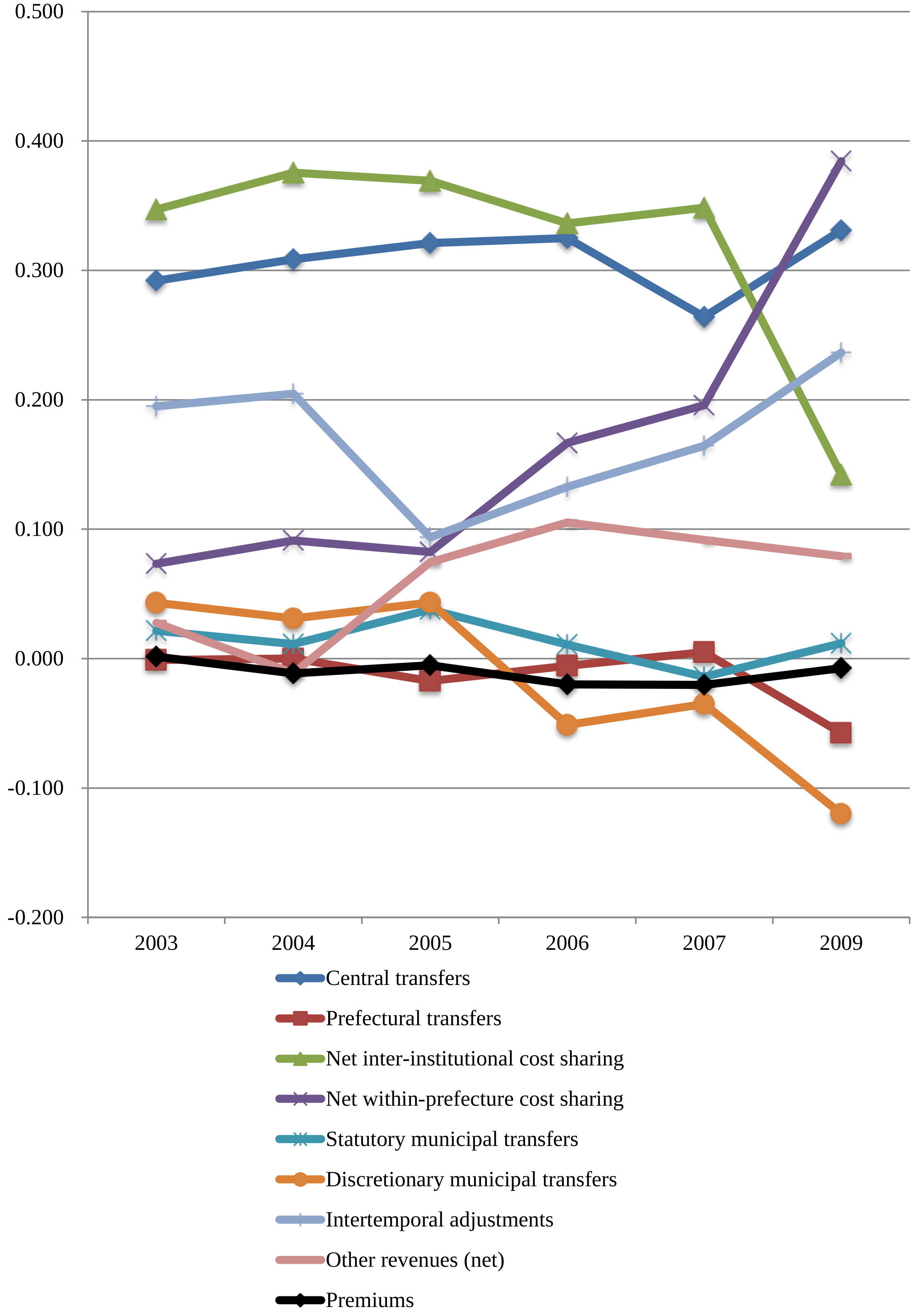

Table 1. Public health care insurance in Japan

<table>
<tr><th colspan="3">Institutional Types</th><th>Insurer/Managing Organization</th><th>Population Covered</th></tr>
<tr><td rowspan="7">Employees' Health Insurance (EHI)</td><td colspan="2">Japan Health Insurance Association (JHIA)-managed</td><td>JHIA</td><td>34 million</td></tr>
<tr><td colspan="2">Association-managed</td><td>EHI Associations</td><td>28 million</td></tr>
<tr><td colspan="2">Seamen's Health Insurance</td><td>JHIA</td><td>.16 million</td></tr>
<tr><td rowspan="3">Mutual Aid Associations</td><td>central government employees</td><td>Central Government Employees' Mutual Aid Associations</td><td rowspan="3">9 million</td></tr>
<tr><td>local government employees</td><td>Local Government Employees' Mutual Aid Associations</td></tr>
<tr><td>private school employees</td><td>Private School Teachers and Employees Association</td></tr>
<tr><td rowspan="2">National Health Insurance (NHI)</td><td colspan="2">Municipal NHI</td><td>Municipalities</td><td rowspan="2">42 million</td></tr>
<tr><td colspan="2">NHI Association-managed</td><td>NHI Associations</td></tr>
<tr><td colspan="3">Health Care Service for the Old</td><td>prefecture-based large area unions</td><td>13 million</td></tr>
</table>

Source: Author's construction based on various government documents.

Table 2. Shares of items in the National Health Insurance (NHI) special accounts

	2002	2003	2004	2006	2007	2008	2009
Premiums	60.8%	52.6%	50.1%	47.3%	44.1%	35.9%	34.9%
Central transfers	61.0%	56.2%	53.3%	45.5%	38.3%	35.8%	36.5%
Provincial transfers	0.4%	1.0%	1.0%	5.6%	6.7%	6.5%	6.5%
Net transfer from inter-institutional cost sharing	−40.5%	−27.6%	−19.6%	−13.2%	−3.1%	11.1%	11.5%
Net transfers from within prefectural cost sharing	0.7%	0.0%	0.0%	0.0%	0.0%	0.0%	0.0%
Statutory municipal transfers	11.0%	11.0%	10.3%	9.8%	9.3%	8.0%	8.0%
Discretionary municipal transfers	6.6%	5.9%	5.5%	5.1%	4.4%	4.3%	4.1%
Inter-temporal adjustment	3.48%	3.15%	1.86%	1.80%	0.75%	-0.69%	-0.40%
Other revenues (net)	−1.5%	−0.7%	−0.7%	−0.8%	−1.2%	−1.4%	−1.4%

Source: Author's calculation based on data obtained from the Ministry of Health, Labour and Welfare (MHLW).

Table 3 Major institutional changes in municipal National Health Insurance (NHI) in the 2000s

FY 2002	• The matching rate of central transfers for catastrophic medical expenses has changed to the current 25%. • Copayments for the insured aged less than 3 years were increased to 20% of actual medical fees. • The Ministry of Health, Labour and Welfare (MHLW) changed the medical fee system.
FY 2003	• The ceiling of catastrophic medical expenditures to which both central and prefectural transfers are paid was reduced from JPN¥800,000 to JPN¥700,000, resulting in a decrease in such transfers.
FY 2004	• The MHLW changed the medical fee system.
FY 2005	• The matching rate of central transfers for medical benefits decreased from 40% to 36%. • The aggregate coverage of the Central Adjustment Grant (a central transfer) for medical benefits was reduced from 10% to 9%. • The Prefectural Adjustment Grants was introduced. • The matching rate of a prefectural grant for one type of municipal transfer to the NHI special account was increased from 25% to 75%.
FY 2006	• The central matching rate for medical benefits decreased from 36% to 34%. • The matching rate for the Prefectural Adjustment Grants increased from 5% to 7%. • The medical benefits for hospital stays and meals were reduced. • The co-payments for the elderly with income above a certain level were increased from 20% to 30%. • The within-prefecture cost sharing scheme for catastrophic medical expenditures more than JPN¥300,000 was introduced. • The MHLW changed the medical fee system.
FY 2008	• Those aged 75 years and above who had been enrolled in the NHI became covered by the Health Care Service for the Old-Old (HCSOO), where each NHI insurer contributes funds to the HCSOO to finance the medical benefits of the "old-old" through a new system of inter-institutional transfers. • A new equalizing scheme for the expenses of those aged between 65 and 75years in social insurance schemes (including the NHI) was introduced (inter-institutional transfers). • The MHLW changed the medical fee system.

Source: Author's descriptions based on various government documents.

Table 4. Sample statistics

	FY 2002–2003 (N = 3129)				FY 2003–2004 (N = 2,470)			
	Mean	S.D.	Min.	Max	Mean	S.D.	Min.	Max.
Medical benefits	16,820	9,039	-76,423	100,984	8,891	8,492	-69,898	70,307
Central transfers	2,051	11,317	-111,839	140,878	922	8,511	-123,200	73,130
Provincial transfers	974	461	-9,540	10,934	-10	609	-9,893	17,225
Inter-institutional cost sharing (net)	12,151	12,407	-82,561	129,337	8,916	10,321	-79,559	102,396
Within-prefectural cost sharing (net)	-816	1,755	-18,778	21,304	23	1,874	-15,539	17,313
Statutory municipal transfers	1,600	2,983	-49,900	30,603	-294	2,525	-28,126	19,046
Discretionary municipal transfers	106	4,139	-51,184	53,291	301	5,851	-41,939	185,591
Inter-temporal adjustment	2,468	10,651	-63,356	105,144	-1,484	10,613	-79,788	102,723
Other revenues (net)	866	9,451	-186,955	132,110	-182	9,271	-259,269	178,213
Premiums	-2,579	3,935	-33,985	17,964	699	4,229	-41,823	30,792
	FY 2004–2005 (N = 1759)				FY 2005–2006 (N = 1,816)			
Medical benefits	11,781	8,621	-44,593	96,532	7,384	8,459	-61,807	120,043
Central transfers	-5,943	8,768	-58,452	190,562	-3,401	8,700	-188,771	145,824
Provincial transfers	7,994	2,599	1,979	51,405	2,943	2,113	-42,798	19,422
Inter-institutional cost sharing (net)	8,063	9,528	-111,279	69,959	7,963	8,879	-57,780	80,704
Within-prefectural cost sharing (net)	7	1,653	-12,405	13,240	79	3,093	-16,910	32,630
Statutory municipal transfers	381	2,416	-34,803	14,743	637	2,449	-25,758	28,833
Discretionary municipal transfers	230	6,951	-82,875	148,168	-354	6,469	-144,155	123,829
Inter-temporal adjustment	870	9,014	-56,725	66,902	-2,059	10,384	-146,640	94,041
Other revenues (net)	-731	12,283	-359,995	172,218	-172	10,099	-61,033	359,255
Premiums	911	3,926	-21,738	21,697	1,747	3,810	-14,828	25,827
	FY 2006–2007 (N = 1803)				FY 2007–2008 (N = 1,788)			
Medical benefits	14,929	9,547	-48,303	149,326	65,252	25,743	-4,162	233,781
Central transfers	1,287	7,590	-124,252	73,969	18,886	11,126	-36,943	140,073
Provincial transfers	721	1,833	-12,532	33,304	4,132	3,359	-14,104	51,941
Inter-institutional cost sharing (net)	7,194	9,538	-66,031	82,235	37,311	25,987	-70,020	151,197
Within-prefectural cost sharing (net)	27	3,849	-27,627	38,722	-35	6,658	-73,820	81,793
Statutory municipal transfers	573	2,756	-36,873	53,887	2,632	3,804	-36,902	52,975
Discretionary municipal transfers	382	7,721	-245,734	53,544	951	7,001	-54,298	95,183
Inter-temporal adjustment	2,958	11,222	-58,354	166,553	-4,860	15,985	-249,317	101,237
Other revenues (net)	42	6,481	-41,262	139,936	-2,042	5,781	-80,086	41,035
Premiums	1,746	3,089	-13,126	22,077	8,276	6,774	-10,684	36,906
	FY 2008–2009 (N = 1,721)				All observations (N = 14,486)			
Medical benefits	6,480	13,545	-140,391	88,851	18,187	22,104	-140,391	233,781
Central transfers	1,691	11,205	-96,624	63,630	2,144	11,955	-188,771	190,562
Provincial transfers	877	4,079	-44,744	43,722	2,252	3,394	-44,744	51,941
Inter-institutional cost sharing (net)	2,314	15,467	-101,665	114,374	11,898	17,146	-111,279	151,197
Within-prefectural cost sharing (net)	-1,252	8,347	-82,864	49,506	-311	4,314	-82,864	81,793
Statutory municipal transfers	676	3,140	-21,834	53,237	898	3,025	-49,900	53,887
Discretionary municipal transfers	891	8,002	-58,738	149,680	329	6,470	-245,734	185,591
Inter-temporal adjustment	2,308	17,901	-153,170	132,072	170	12,619	-249,317	166,553
Other revenues (net)	-801	6,464	-101,696	45,816	-296	8,949	-359,995	359,255
Premiums	-224	4,091	-21,087	23,847	1,104	5,343	-41,823	36,906

Notes: All variables are differenced, measured in JPN¥, and in per subscriber terms.

Table 5. Degree of risk sharing (estimation results)

	FY 2003	FY 2004	FY 2005	FY 2006	FY 2007	FY 2008	FY 2009
Premiums	0.2 (0.9)	–1.1 (1.0)	–0.5 (1.2)	–2.0 (1.2)	–2.0** (1.0)	3.8*** (0.7)	–0.7 (0.9)
Central transfers	29.2*** (3.6)	30.9*** (3.0)	32.1*** (3.7)	32.5*** (4.8)	26.4*** (8.2)	20.0*** (1.7)	33.1*** (3.3)
Provincial transfers	–0.1 (0.2)	0.0 (0.2)	–1.7 (2.5)	–0.6 (1.3)	0.5 (0.9)	4.5*** (0.3)	–5.8*** (2.2)
Net transfers from inter-institutional cost sharing	34.7*** (4.7)	37.6*** (4.3)	36.9*** (4.8)	33.6*** (4.5)	34.8*** (4.3)	68.1*** (2.7)	14.2** (5.8)
Net transfers from within prefectural cost sharing	7.3*** (0.9)	9.1*** (0.9)	8.2*** (0.9)	16.7*** (2.0)	19.6*** (2.5)	4.8*** (1.2)	38.4*** (2.8)
Statutory municipal transfers	2.1** (1.0)	1.1 (0.7)	3.8*** (1.3)	1.1 (1.2)	–1.4 (2.3)	3.6*** (0.5)	1.2 (0.8)
Discretionary municipal transfers	4.3*** (1.4)	3.1 (2.2)	4.3 (6.1)	–5.1 (3.5)	–3.5 (5.5)	–2.0*** (0.8)	–12.0* (6.8)
Inter-temporal adjustment	19.5*** (3.1)	20.5*** (4.0)	9.4** (4.2)	13.3** (6.1)	16.4* (9.8)	0.0 (2.1)	23.6*** (5.7)
Other revenues (net)	2.8 (3.3)	–1.1 (2.6)	7.4 (10.0)	10.5 (7.8)	9.2** (4.5)	–2.7*** (0.7)	7.9** (4.0)

Notes: *** $p \leq .01$, ** $.01 < p \leq .05$; * $.05 < p \leq 0.10$. The unit is in percentage. The robust standard errors are in parentheses.