Physicians' Responses to Medical Subsidy Programs*

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Abstract

Previous studies have found that expanding health insurance enrollment encourages physicians to work in underserved areas where many uninsured lived. However, in countries with universal health coverage, a further expansion of the generosity of health insurance may not have such an equalization effect. Instead, it may incentivize physicians to operate in urban areas, generating a concentration of physicians in cities. To test this hypothesis, we examine how the large expansion of a medical subsidy program changes the behavior of primary care physicians. In Japan, a local subsidization program, which significantly reduces out-of-pocket expenses for children's healthcare utilization, referred to as the Medical Subsidy for Children and Infants (MSCI), has rapidly spread across the country since 2000, although there are significant regional differences in eligibility criteria. By using a census of clinics from 1999 to 2011, matched with municipality-level eligibility criteria of the MSCI, we implement difference-in-differences-in-difference analysis. The results show that MSCI increases the monthly number of visits per clinic. However, physicians choose to operate in more densely populated areas, suggesting that expanding the generosity of the health insurance system may accelerate the concentration of physicians into cities.

Keywords : Location choice; Natural experiment; Primary care; Children; Extensive and intensive margin

JEL classification : J13, J16, I10

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1 Introduction

Since a seminal paper by Fuchs (1978), health economists have repeatedly investigated the consequences of excessive concentration of physicians in local healthcare markets. Like any other professionals, physicians choose to live in places they like to provide healthcare. Thus, their choice of location depends not only on financial considerations but also their leisure time preferences. For example, physicians who love fashion may choose to operate in metropolitan areas even with the high competition of such areas. On the other hand, those who love the mountains may choose to live in rural areas where the number of patients they treat is limited. Eventually, the spatial distribution of physicians is not compatible with an overall efficient and equitable healthcare system.

This point is particularly important in light of the objective of the health system to provide quality healthcare services with affordable fees. In fact, physician shortages and the uneven distribution of physicians have been regarded as important policy issues (Newhouse, 1990; Anand *et al.*, 2011; Tanihara *et al.*, 2011). Policy-makers have struggled to solve this problem over the last several decades, but uneven distribution persists. An international comparison study indicates that physicians are distributed unequally across different regions in virtually all OECD countries (Ono *et al.*, 2014).[1] Many studies support these political endeavors because an excess concentration of physicians in urban areas may lead to abuses of healthcare resources (Fuchs, 1978) and insured persons in rural and remote areas have very limited access to healthcare services (Lu & Slusky, 2016).[2]

A recent epidemiological systematic review proposes one solution to this problem. Grobler *et al.* (2015) reviewed 98 full-text articles on the impacts of policy interventions targeting the uneven distribution of physicians and concluded that the expansion of health insurance coverage may alleviate the concentration of physicians in urban areas (Yang *et al.*, 2013). Similar findings are presented in more recent economic studies by Chen *et al.* (2017) and Huh (2017). They reveal that the expansion of children's health insurance and dental insurance in the US may lead medical professionals to work in underserved and poor areas because of the underlying demand for healthcare in these areas.[3] Thus, expanding *extensive margin* health insurance coverage may have favorable consequences in terms of the geographical distribution of physicians.

However, this conclusion does not hold when we consider the impacts of health insurance expansion in its *intensive margin*, which means the generosity of insurance coverage. In the case of intervention on the extensive margin, already-insured persons are mostly unaffected.[4] Thus, healthcare demand increases mainly in poor and rural areas where many uninsured live (Yang *et al.*, 2013; Chen *et al.*, 2017; Huh, 2017). By contrast, the expansion of the intensive margin affects all insured persons and increases healthcare

[1] According to the 2012–13 OECD Health System Characteristics Survey, only one out of 34 OECD countries does not consider the distribution of doctors to be an issue (namely, the Netherlands).

[2] In the US, 20% of the population lives in rural areas, where they represent one of the largest medically underserved populations (Rabinowitz *et al.*, 1999).

[3] Chen *et al.* (2017) investigated the reauthorization of Child Health Insurance Programs in 2009. Using the data on physicians in training in New York, they found that newly trained pediatricians are more likely to enter private practice. In addition, they found evidence suggesting increased physician supply in rural areas. Huh (2017) also found that dentists were more likely to choose rural areas as their residence after the expansion of adult Medicaid dental coverage in the US.

[4] Some move to expanded public insurance from private insurance (Cutler & Gruber, 1996)

demand not only in rural areas but also in urban ones. Therefore, the expansion of the intensive margin does not alleviate physicians' concentration, and worse still, it would theoretically exacerbate the concentration of physicians in urban areas.[5] The distinction between the extensive and intensive margin of health insurance is particularly important because many developed countries, where universal health coverage has been already achieved, have also reduced the share of out-of-pocket expenses (Fan & Savedoff, 2014; Dieleman *et al.*, 2017). Therefore, this study examines how the large-scale reduction of out-of-pocket charges impacts primary care physician labor supply and practice location choice.

As a conceptual framework, we adapt the core periphery model (the so-called new economic geography model; Krugman (1991)) by replacing the migration of skilled workers with the migration of physicians. Over the last three decades, the core-periphery model has been a standard model for dealing with the geographic maldistribution of economic activity across regions[6], because this model is based on the micro-founded models consistent with individual optimization and market clearing. The agglomeration forces in the core-periphery model mainly arise from a preference for variety. Workers who prefer variety are attracted to urban core areas that have many varieties of goods and services compared with rural periphery areas. Physicians would be also attracted to urban areas because of their preference for variety, but an increase in physicians in urban areas decreases their profits (Newhouse *et al.*, 1982).[7] Assuming that physicians prefer variety, we theoretically show that reduced patient cost-sharing strengthens physician concentration in urban areas.

To gain empirical evidence on the effects of reduced cost-sharing, recent experiences in Japan may provide a valuable opportunity because there has been a large-scale expansion of subsidization programs for out-of-pocket costs implemented by local governments. In Japan, the national government sets coinsurance for preschool children at 20% countrywide, but municipal and prefectural governments can reduce this amount at their own expense. This subsidy program, referred to as the Medical Subsidy for Children and Infants (MSCI, in Japanese: *Nyuyoji Iryohi Jyosei*), has been dramatically expanded in the last two decades. Because the number of municipalities in Japan exceeds 1700 and each municipality sets different eligibility ages for subsidy entitlement, (up to 22 years old) these variations in eligibility age could be exploited during the 2000s.

These extensive regional diversities in the MSCI eligible ages serve as plausible quasi-experiments that may uncover the market-wide effects of reduced cost-sharing. In empirical analysis, the municipality-level eligibility criteria obtained from Takaku (2016) are matched with all clinics around Japan in service between 1999 and 2011. Among the full sample of pediatric clinics (N = 90,703), 62,221 are included in our analysis. In primal empirical strategy, we apply difference-in-differences type estimation by exploiting a geographically staggered expansion of MSCI eligible age as a natural experiment. In addition, following the empirical strategy used by Garthwaite (2012), we compare clinics that provide care mainly for children

[5]In our theoretical model presented in Section 3, we suggest that the expansion of the intensive margin induces physicians to choose an urban location, assuming that their utility function incorporates a preference for a variety of goods.

[6]Krugman & Venables (1995), Venables (1996), Puga (1999), and Helpman (1998) are representative studies engaging in this model. For earlier reviews of the literature, see for example Overman *et al.* (2003) and Redding (2011).

[7]In Japan, several studies also found that physicians' wages increase if they work in rural areas (Sano & Kishida, 2004; Ikegami, 2014)

(N = 12,180, "child clinic") and those that provide care for both children and adults (N = 55,877, "all-generation clinic") because the MSCI expansion would have affected the former clinics more than the latter.

Three major findings are presented. First, the results show that the extension of the MSCI has greatly increased the number of visits at clinics that mainly treat children (*child clinics*). Numerically, the extension of the MSCI eligibility age by 10 years increases the total number of visits at child clinics by approximately 5.9%. This quantitative impact is slightly smaller than the price elasticity of health care utilization presented by the RAND Health Insurance Experiment (Manning *et al.*, 1987), but is fairly consistent with a recent study that investigated the effects of the MSCI on patient-level data (Iizuka & Shigeoka, 2018). Second, newly established child clinics are more likely to practice in areas with high population density under the generous MSCI system, supporting the hypothesis that generous health insurance accelerates the uneven distribution of physicians. Third, we find a statistically significant decrease in the number of consultation days among child clinics. While this finding is consistent with Enterline *et al.* (1973) and Garthwaite (2012) who show that health insurance expansion leads to less working hours among primary care physicians,[8] the quantitative impact on this outcome is very small and almost negligible.

This paper has two contributions to the literature on the supply-side effects of health insurance expansion (Finkelstein, 2007; Kondo & Shigeoka, 2013). First, we construct more accurate continuous measures for the characteristics of practice location. Even though some studies deal with similar issues, (Escarce *et al.*, 1998; Polsky *et al.*, 2000; Yang *et al.*, 2013; Chen *et al.*, 2017),[9] their measurement of practice location was based on binary information such as rural vs urban. In the present study, we calculate the population density within a radius of a few kilometers for all clinics from the exact address of all clinics and mesh population data from the Census. This density is regarded as a standard measure of the location's urbanity. Second, by using more accurate measures, we reach the opposite conclusion with previous studies that investigated the effects of health insurance expansion in the extensive margin (Yang *et al.*, 2013; Grobler *et al.*, 2015; Chen *et al.*, 2017). As is suggested by our theoretical model, it is possible that a large-scale expansion of health insurance in the intensive margin may accelerate the concentration of primary care physicians into urban areas. If so, abolishing the provision of free healthcare is far from a sufficient policy to provide access to health services in underserved areas because these areas become more unlikely to be chosen as physicians' practice location. Considering all of the above, this study suggests different effects of health insurance expansion between the expansion margins.

The paper is constructed as follows. Section 2 provides institutional background of the study. Section

[8] Enterline *et al.* (1973) interviewed Canadian practicing physicians before and after the introduction of a universal healthcare program in Quebec, and found that physicians decreased their hours worked following the implementation of a universal healthcare program. They conclude that with the advent of free medical care, physicians may now question the necessity of working 50 or 60 hours a week. Garthwaite (2012) also found the expansion of Children's Health Insurance in the US reduced pediatricians' hours worked.

[9] Among these studies, Escarce *et al.* (1998) and Polsky *et al.* (2000) investigated the effects of the penetration of health maintenance organization (HMO) in the US. Escarce *et al.* (1998) investigated how new physicians' location choices was associated with HMO penetration, whereas Polsky *et al.* (2000) focused on practicing physicians' location choices. Importantly, Polsky *et al.* (2000) found minor effects of HMO penetration on existing physicians' relocation and geographical redistribution, whereas Escarce *et al.* (1998) showed that new physicians' location choice was responsive to HMO penetration.

3 explains the theoretical model that motivated our empirical analysis. Section 4 explains the data. In section 5, we describe our empirical methodology, with the main results reported in section 6. In section 7, we implement a variety of robustness checks. Heterogeneity of the MSCI effects is examined in section 8. Finally, section 9 provides concluding remarks.

2 Background

2.1 Japan's Healthcare System

Japan introduced universal health coverage in 1961(Ikegami & Campbell, 1995; Kondo & Shigeoka, 2013). Enrollment in health insurance is mandatory. Out-of-pocket expense for healthcare utilization is calculated as a coinsurance, which is a percentage of total healthcare costs. This system is applied to all patients regardless of their ages. The national-level coinsurance is 30% for school-age children, which is the same rate for adult patients, and it is slightly reduced to 20% for preschool-age children. Of course, the expenses implied by these coinsurances are not especially low compared with other developed countries. In fact, among the G7 countries, children's healthcare is basically free in the UK, Germany, France, Italy and Canada. Some of these countries require out-of-pocket expenses for adult patients, but children are treated for free. From the international perspective, the US and Japan, where children's healthcare is not free, are two exceptions. Unlike the US, however, out-of-pocket expenses do not differ across insurance plans in Japan. Even though the Japanese health insurance system is highly fragmented with more than 1500 insurers, coverage for healthcare services and out-of-pocket expenses are the same[10] and adverse selection is not relevant.

To access primary care physicians, Japan has a unique feature referred to as the "free access system". This means that there are no gatekeepers, such as in the UK, and all patients can freely choose physicians when they receive treatment. Note that Japan also differs from the US system where insurers have considerable power over the choice of medical providers though managed care. Probably due to the large freedom of choice of medical providers, the use of healthcare services has not severely constrained, regardless of high out-of-pocket costs. According to the Organisation for Economic Co-operation and Development (OECD) health statistics, the number of doctor consultations per capita in Japan has continuously been one of the highest since the 1980s.

When we look at the institutional differences between Japan and the US more deeply, one important point that becomes apparent is that the medical costs for primary care in Japan are based on fee-for-service and calculated upon a single reimbursement scheme (Ikegami & Campbell, 1995). Thus, physicians receive the same reimbursement for the same treatment anywhere in Japan. For example, the basic reimbursement for clinics is generally 3,500 JPY for a patient with a common cold regardless of their insurance plans and residential location[11]. Given that the effects of the public health insurance expansion are usually discussed on the mixed-economy model, which postulates a differential reimbursement rate across insurance plans

[10] Premium rate differs across insurance plans

[11] First visit fee is 2,820 JPY and prescription fee is 680 JPY.

(Sloan *et al.*, 1978; Garthwaite, 2012),[12] a unified reimbursement scheme in Japan makes the interpretation of our analysis more straightforward.

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2.2 Medical Subsidy for Children and Infants

Regardless of the high utilization rate of outpatient care in Japan, out-of-pocket expenses for children's healthcare utilization has fallen dramatically over the last two decades. Both the declining birth rate and increasing child poverty rate are contributing factors to this phenomenon. Because healthcare costs for childhood illness are sometimes high, a reduction in coinsurances through the appropriation of local tax revenue has been strongly supported in the local political environment. Therefore, municipalities began the supplemental subsidization program, namely the MSCI. Note that Japan has been divided into 1748 municipalities as of 2012. Thus, the geographically staggered expansion of MSCI provides a plausible natural experiment in such a homogeneous population.

Eligibility for MSCI is mostly determined by child age. Municipalities can freely set the maximum age for MSCI eligibility under which any child can receive subsidy. This maximum age, which is referred to as "eligible age", grew gradually over the last two decades. While eligible age is the most important data point in our analysis, the MSCI system differs across municipalities in three other aspects. First, municipalities can restrict eligibility of the MSCI for children by excluding children from high-income households. This ceiling is adopted by 34% of all municipalities. Second, municipalities can also choose the reimbursement method (in-kind transfer or refund). Finally, the amount of subsidy varies across municipalities. Some municipalities charge very small out-of-pocket expenses to reduce moral-hazard, while the copayment is completely waived in 54% of all municipalities (Ministry of Health & Welfare, 2013).

3 Theoretical Framework

To motivate the empirical work that follows, we set up a conceptual framework that links health insurance expansion and physician's practice location choice, by modifying a core-periphery model (Krugman, 1991).[13] Although many factors may be potentially associated with practice location choice, including hospital location, demographics, financial factors, mal-practice insurance premiums, and health insurance expansion for uninsured population (Newhouse *et al.*, 1982; Frank, 1985; Polsky *et al.*, 2002; Chou & Lo Sasso, 2009; Aiura, 2011; Chen *et al.*, 2017), the association between the health insurance system and practice location choice is not so obvious. The present model intends to show how the effect of health insurance expansion on physicians' practice location differs between the extensive and intensive margin of health insurance.

[12] In the mixed-economy model proposed by Sloan *et al.* (1978), the reimbursement schedule for physicians should be non-linear because physicians face two markets. For the privately insured patients, the demand curves faced by the physicians are downward sloping, but reimbursement is fixed in public health insurance.

[13] Aiura (2011) analyzed the geographic maldistribution of physicians across regions with a modified core-periphery model. Whereas Aiura (2011) considers specialist physicians and assumes that their services are horizontally differentiated, the present model considers primary care physicians and assumes that their services are homogeneous.

3.1 Model

We constructed an economic model that consists of multiple areas: area 1, area 2, $\cdots$, area k, $\cdots$. The population is different in each area, and the population in area k is denoted by N_k. Here, the labor population share in area k is denoted by δ_k and we assume that labor is fully employed. Thus, the population of workers is equal to $\delta_k N_k$. Moreover, we assume that the labor population share in a high-population area is more or equal to that of low-population areas (i.e., $\delta_k \geq \delta_l$ if $N_k > N_l$), because the former would be likely to attract labor force.[14] The economy in each area has three sectors: the tradable sector, the non-tradable sector, and the healthcare sector. The tradable sector is perfectly competitive and produces homogeneous goods, which is costlessly traded between areas, under constant returns using labor as the only input. The non-tradable sector is monopolistically competitive and produces a continuum of a variety of horizontally differentiated products and services under increasing returns using labor as the only input (Chamberlin, 1933).[15] Labor is mobile between tradable and non-tradable sectors. In the healthcare sector, self-employed physicians provide homogeneous service. We assume that only self-employed physicians can choose the area of their practice (in other words, labor is immobile between the areas) because we focus on physician's practice location choice. [16]

Because the output of the tradable sector is homogeneous and traded without transportation costs, it is chosen as numéraire with price normalized to 1. Under free entry, the wages of workers are the same as the price of goods in the tradable sector, and equal to 1. Because labor is mobile between tradable and non-tradable sectors, the wages of workers in non-tradable sectors are also equal to 1. Workers pay taxes for social security and public health insurance schemes. The disposable income after deduction of the taxes is denoted by w. Persons not in the labor force receive income, B, which is less than w, from social security benefits.

Preferences are identical across all people and described by a Cobb-Douglas utility,

$$U = Q^{\alpha} H^{\beta} T^{1-\alpha-\beta}, \tag{1}$$

where

$$Q \equiv \left(\int_0^M q(i)^{(\sigma-1)/\sigma} dk \right)^{\sigma/(\sigma-1)}, \quad 0 < \alpha < 1, \quad 0 < \beta < 1, \quad \sigma > 1. \tag{2}$$

$q(i)$ represents the consumption of variety $i \in [0, M]$. H and T represents the consumption of services and goods in the healthcare and tradable sectors, respectively. In this utility, the elasticity of substitution

[14]Population ages 15–64 (% of total) in rural areas is 60.0% in Japan (2010), 65.9% in the US (2010), and 63.1% in the UK (2011), whereas populations aged 15–64 (% of total) in urban areas is 63.6% in Japan (2010), 67.4% in the US (2010), 66.7% in UK (2011) (Source: United Nations Statistics Division).

[15]Even if the products in this sector are tradable at positive costs, the lemma and the proposition in this section remain true.

[16]Because the ratio of physicians to population is very small in typical countries, we assume that the number of physicians is not included in population.

between any two varieties is identical to σ. The budget constraint is given as

$$Y = \int_0^M p(i)q(i)di + T + \Theta p_h H \tag{3}$$

where Y represents the consumer income, and $p(i)$ and p_h represent the price of variety i and healthcare services, respectively. The costs of healthcare services are partially reimbursed in the public health insurance scheme, so out-of-pocket expense of healthcare services is $\Theta p_h H$ where $\Theta \in (0, 1]$ represents coinsurance. The public health insurance covers workers only, whose coinsurance, Θ, is equal to $\theta \in (0, 1)$. We consider two cases regarding whether non-workers are covered by public health insurance. In the first case, we assume that the public health insurance covers non-workers at the same level as the worker. In the second case, we assume that the public health insurance does not cover non-workers, so the coinsurance of them, Θ, is equal to 1. We call the former case the full-coverage case and the latter case the partial-coverage case. By maximizing utility (1) subject to the consumer budget constraint (3), the individual demand functions are given as

$$q(i) = \alpha \frac{P^{\sigma-1}}{p(i)^{\sigma}} Y, \tag{4}$$

$$Q = \frac{\alpha Y}{P}, \tag{5}$$

$$T = (1 - \alpha - \beta)Y, \tag{6}$$

$$H = \frac{\beta Y}{\Theta p_H}, \tag{7}$$

where

$$P \equiv \left(\int_0^M p(k)^{1-\sigma} dk \right)^{1/(1-\sigma)}. \tag{8}$$

Substituting (5), (6), and (7) into (1), we obtain the indirect utility function:

$$V = \left(\frac{\alpha}{P}\right)^{\alpha} \left(\frac{\beta}{\Theta p_H}\right)^{\beta} (1 - \alpha - \beta)^{1-\alpha-\beta} Y. \tag{9}$$

Because the income of workers and non-workers is w and B, respectively, if the number of variety in area k is M_k, the total demand of the firm producing variety i in area k is the follows.[17]

$$q_k(i) = \alpha \frac{P_k^{\sigma-1}}{p_k(i)^{\sigma}} [\delta_k w + (1 - \delta_k)B] N_k, \tag{10}$$

where

$$P_k \equiv \left(\int_0^{M_k} p_k(i)^{1-\sigma} di \right)^{1/(1-\sigma)}. \tag{11}$$

Moreover, if we assume that the firm producing quantity q in the non-tradable sector needs $n + n_1 q$ units

[17] Because there are increasing returns in the non-tradable sector, each variety is produced by a single firm if there are no scope economies.

of labor input, the profit of the firm producing variety i in area k is as follows.

$$\pi_k(i) = (p_k(i) - n_1)q_k(i) - n. \tag{12}$$

Because the non-tradable sector is monopolistically competitive, each firm in the non-tradable sector has a negligible impact on the market in accordance with Chamberlin (1933); that is, it neglects the impact of a price change over P_k. Solving the first-order condition of (12) using (10) yields the equilibrium price.

$$p_k^*(i) = \frac{\sigma}{\sigma - 1} n_1. \tag{13}$$

Note that the equilibrium price is the same for any firms in area k. Substituting (13) into (11), we obtain

$$P_k^* = \frac{\sigma}{\sigma - 1} n_1 \left(\frac{1}{M_k}\right)^{1/(\sigma-1)}. \tag{14}$$

Substituting (13) into (10) and using (12) and (14), we gain the equilibrium profit:

$$\pi_k^*(i) = \frac{\alpha}{\sigma} \cdot \frac{[\delta_k w + (1 - \delta_k)B]N_k}{M_k} - n. \tag{15}$$

Under free entry, firms enter or exit the non-tradable sector unless their profit is 0. Therefore, the variety in area k in the equilibrium is

$$M_k^* = \frac{\alpha}{\sigma n}[\delta_k w + (1 - \delta_k)B]N_k, \tag{16}$$

which implies that the larger the population in the area is, the larger the variety of non-tradable goods. From (9) and (14), the utility increases with an increase in M_k. It implies that people love the variety of non-tradable goods. Accordingly, the larger the population of an area, the higher utility of consumers in the area.

Because individual demand of healthcare services is presented by (7) and the total income in area k is $[\delta_k w + (1 - \delta_k)B]N_k$, the total demand of healthcare services in area k is

$$H_k = \bar{H}_k = \frac{\beta}{\theta p_h}[\delta_k w + (1 - \delta_k)B]N_k, \tag{17}$$

for the full-coverage case (when public health insurance covers all people) and

$$H_k = \hat{H}_k = \frac{\beta}{\theta p_h}[\delta_k w + \theta(1 - \delta_k)B]N_k, \tag{18}$$

for the partial-coverage case (when public health insurance covers only the workers). The total demand of healthcare services would be divided equally among the physicians in the areas, because self-employed physicians provide homogeneous services. Therefore, the income of self-employed physicians in area k is

$$Y_k = p_h \frac{H_k}{D_k} - c \tag{19}$$

where D_k represents the number of physicians in area k and c represents the opportunity costs of becoming a physician.[18] Because the preferences of self-employed physicians are identical to (1), substituting (19) into (9) yields the utility of physicians in area k

$$V_k = \left(\frac{\alpha}{P_k^*}\right)^{\alpha}\left(\frac{\beta}{\Theta p_h}\right)^{\beta}(1-\alpha-\beta)^{1-\alpha-\beta}\left(p_h\frac{H_k}{D_k}-c\right). \tag{20}$$

In the equilibrium, the utility of physicians is the same between areas ($V_k = V_l, k \neq l$), because physicians would practice in areas where they gain higher utility. Therefore,

$$\left(\frac{1}{P_k^*}\right)^{\alpha}\left(p_h\frac{H_k}{D_k}-c\right) = \left(\frac{1}{P_l^*}\right)^{\alpha}\left(p_h\frac{H_l}{D_l}-c\right), \text{ for any } k \text{ and } l, \tag{21}$$

which leads to the following lemma.

Lemma 1. In both full-coverage and partial-coverage cases, as the population increases, the number of physicians per person also increases in the area. (i.e., $D_k/N_k > D_l/N_l$ if $N_k > N_l$).

Proof: See Appendix F.1.

The interpretation of the result is simple. (14) and (20) show that the utility of physicians depends on the variety of non-tradable goods and the physicians' income. If income is the same between areas, physicians prefer an area with a greater variety of non-tradable goods. Because areas with larger populations have a greater variety of non-tradable goods (as (14) shows), physicians concentrate in areas with large populations.

Furthermore, this implies that physician income is higher in low-population areas than in high-population areas[19] and if the expansion of public health insurance increases income disparity between low-population and high-population areas, it improves the uneven geographical distribution of physicians; otherwise, it worsens.

3.2 Expansion of the Extensive Margin of Public Health Insurance

In the partial-coverage case, public health insurance covers only workers, whereas, in the full-coverage case, public health insurance covers all people. Therefore, the change in public health insurance scheme from partial-coverage to the full-coverage expands the extensive margin of public health insurance. The following proposition provides the effect of this expansion on geographical distribution of physicians.

Proposition 1. The expansion of public health insurance coverage from workers to the entire population decentralizes the distribution of physicians because it decreases the difference in the number of physicians per person between any two areas.

Proof: See Appendix F.2.

This implies that the expansion of the extensive margin of public health insurance alleviates physician's

[18] For simplicity, we ignore the other costs. However, the results are not unchanged by it, if the cost function is the same between areas. Further, even if the fixed costs in the area are in proportion to its population size, all lemma and propositions remain true.

[19] This is because we consider primary care in the model. If we consider agglomeration effects or synergy effects among physicians having different skills, physicians' incomes may be higher in high-population areas than in low-population areas. However, such effects would rarely occur because primary care is standardized and homogeneous.

concentration into high-population areas, which is consistent with the results of Yang *et al.* (2013), Chen *et al.* (2017) and Huh (2017). We interpret the result accordingly. Health insurance expansion increases the utilization of healthcare services in low-population areas more than in high-population areas because the share of the uninsured population generally increases as the population increases in an area. Also, our model suggests that, at the market equilibrium, the number of physicians per capita is higher in high-population areas than in low-population areas. This indicates that physicians in lower population areas experience a greater surge in healthcare demand when health insurance is expanded in the extensive margin. As a possible result of these effects, the expansion of the extensive margin increases physicians' income levels in low-population areas more than in high-population areas, which encourages physicians to open their clinics in low-population areas.

3.3 Expansion of the intensive margin of public health insurance

A decline in the coinsurance rates increases the healthcare demand of each person who joins public insurance (see (7)) and thus expands the intensive margin of public health insurance. The following proposition provides the effect of this expansion on the geographical distribution of physicians.

Proposition 2. In both full-coverage and partial-coverage cases, a decrease in coinsurance strengthens the concentration of physicians in areas with large populations, because it increases the difference in number of physicians per person between any two areas. Especially, in full-coverage case, even if we do not assume that the labor population share in a high-population area is more or equal to that of low-population areas, a decrease in coinsurance strengthens the concentration of physicians in areas with large total populations, not with large labor populations.

Proof: See Appendix F.3

This proposition implies that the expansion of the intensive margin of public health insurance accelerates the concentration of physicians into high-population areas, which is the opposite to the result for Proposition 1. We interpret the result of Proposition 2 accordingly. The expansion of the intensive margin increases the revenue of each physician at the same rate in all areas. On the other hand, the fixed opportunity costs of becoming a physician is unchanged by the expansion and are the same between areas. Therefore, the income disparity of physicians decreases between areas because of the expansion. Under the utility function in the model, people prefer areas with a greater variety of non-tradable goods even if they receive lower real incomes in this area. Accordingly, when health insurance expands its coverage, the income loss associated with choosing high-population areas instead of low-population areas is reduced, which then exacerbates the uneven geographical distribution of physicians. Furthermore, Proposition 2 implies that when health insurance expands its coverage in full-coverage case, physicians are attracted by total population regardless of population of workers. For example we assume the case that the total population is a bit larger in area j than in area k but the labor population is sufficiently larger in area k than in area j. In this case, an increase in the demand for healthcare with an expansion of the intensive margin is larger in area k than in area j because workers more increase the demand for healthcare than non-workers do. In full coverage case, however, the increasing rate of the demand for healthcare in workers is as same as in non-workers, and the growth rate of physicians' revenue is not different between areas. Therefore, an expansion of the intensive margin does not attracts physicians to the area with large demand. When we apply this theory to the recent expansion of MSCI, it is reasonable to predict that the MSCI expansion accelerates the concentration of physicians in high-total-population areas regardless of the children population, because the MSCI expansion is considered as a health insurance expansion in the intensive margin under full-coverage. In the following sections, we examine whether this prediction is supported by the empirical findings.

4 Data

4.1 MSCI Eligibility

Even though institutional setting of the MSCI expansion is unique, the history of the eligible age for the MSCI has not been ascertained from any official surveys, as is noted in Iizuka & Shigeoka (2018). Therefore, Takaku (2016) conducted an original survey for all municipalities in October 2013. Although the response rate was only 55%, 949 municipalities replied to the survey. In addition, after the previous study was published, we implemented a follow-up study to construct more comprehensive data on MSCI eligibility. Specifically, we chose 100 municipalities with the largest populations which did not reply to our previous survey and sent them the same questionnaire. We received responses from 64 municipalities. From 1013 municipalities, we focus on the 614 municipalities in which the MSCI eligibility is consistently available from 1999 to 2011. The reason why some municipalities are dropped is due to the Heisei Great Amalgamation during the mid-2000s. Many amalgamated municipalities could not answer the eligibility criteria before the amalgamation.[20]

Figure 1 presents the municipality-level map of eligible age in 2011. Despite the incomplete response rate, Figure 1 shows that our survey covers a sufficiently large area of Japan and displays large regional disparities in the MSCI eligible age. We summarize the prefecture-level variation of the average eligible age from 1999 to 2011 in Appendix A1. For example, the average eligible age for the MSCI was restricted to under 2 years old in many municipalities in 1999, but it was raised to over 12 years old in many municipalities in central mainland by 2011, while other municipalities still restricted it to preschool children. In an empirical analysis, we use these regional variations of the MSCI eligibility to uncover the causal effects of health insurance expansion on supply-side behaviors. In A2, we also summarize distribution of MSCI eligible age by year.

4.2 Census Data on Clinics

The primary dataset for this analysis is the Survey of Medical Institutions (SMI). This dataset, conducted every 3 years in October, is a national census of hospitals and clinics. The analysis uses the five waves

[20]Though Iizuka & Shigeoka (2018) also hand-collected the detailed information of the MSCI, their MSCI data cover the period from 2005 to 2015. However, even in 2005, MSCI eligible ages were already expanded to cover all preschool age children in 65 % of municipalities in Table A2. Given that healthcare utilization is far higher in preschool-age children than in school-age children, the current study requires information from the period after MSCI was expanded for much younger children to quantify the impacts on clinics rather than patients.

from 1999 to 2011 (1999, 2002, 2005, 2008, and 2011), which cover the entire period when the MSCI was rapidly expanded.[21]

In the SMI data, the official specialties of all clinics are ascertained using a consistent definition. Pediatric clinics are defined as those stating that pediatrics is their specialty. Note that this definition includes clinics where most of the patients are actually adults because many clinics have multiple specialties.[22]

4.3 Dependent Variables and Covariates

On the dependent variables, we first quantify the impacts on the monthly number of visits. This outcome is used to explore how patients, rather than clinics, respond to the MSCI expansion, but it is useful to check the effects because numerous studies have already found that patients are more likely to visit clinics under a generous health insurance system (Iizuka & Shigeoka, 2018). Not finding statistically significant effects on this outcome suggests that our empirical strategy is not particularly credible. In addition to the monthly number of visits, we also investigate how the MSCI expansion affects the number of first visits, follow-up visits, and off-hour visits.

After checking the effects on visits, we explore the impacts on location choice of clinics. A unique feature of the dataset is that the exact address of all clinics is recorded. Using this address information, we investigate how the MSCI expansion is associated with the urbanity of the location of each clinic. Although city amenity encompasses multiple aspects, we calculate the population around clinic location as a proxy for the urbanity of the location. Specifically, we calculate the population density within a radius of 1, 3, and 5 kilometers for all clinics by merging the address of the clinics with population mesh data. In addition, we calculate the population density by the smallest administrative area (SAA) where each clinic is located.[23] Because the number of the SAA is sufficiently large (i.e., 217,186 in the 2010 Census), the population density of SAA may provide accurate information on the urbanity of the location. We use the SAA-based measures in order to supplement the results from the standard measures of the location because population density by age groups is only available in the SAA-based measures. In actual procedure, census results from 2010 are spatially merged with the clinic location by ARC-GIS version 10.3.1. Figure 3 compares the distribution of population density where each clinic is located by 4 difference definitions. In short, population density from the SAA is similar with that of 1 km radius. We also construct a binary variable on the location, based on population with in a 3 kilometer radius. In this variable, clinics are categorized to "rural" if their population density is less than 25 percentile of the entire distribution. This variable is comparable to the measures of the practice location which are used in previous studies (Yang *et al.*, 2013; Chen *et al.*, 2017; Huh, 2017).

Finally, we examine the effects on the number of consultation days per week. Our survey asks, for

[21]Use of census data is also important in the literature of supply-side effects of health insurance because of the findings of some important previous studies' sampling surveys on healthcare providers (Enterline *et al.*, 1973; Baker & Royalty, 2000; Garthwaite, 2012; Buchmueller *et al.*, 2016).

[22]Physicians in Japan can freely choose the official specialty of their clinics from 38 specialties, which is allowed by Medical Basic Law. After reporting and registering their official specialties to the municipality where their clinics are located, physicians can open their clinics.

[23]The SAA (in Japanese: *Cyou Cyou Moku*) divides Japan into 217,186 small areas.

example, whether or not the clinic is open on Mondays. Similar questions are asked about other days of the week. With this information, we count the number of consultation days in a usual week. This variable is not a one-to-one correspondence with working hours investigated in previous studies (Garthwaite, 2012; Buchmueller *et al.*, 2016), but includes useful information about the choice of working hours among physicians. In particular, the number of consultation days has suitable features for our analysis because physicians can choose working hours with complete autonomy, not being fluctuated by the changes in regional healthcare demand.

For the independent variables, we control for various clinic-level covariates, including the number of beds, ownership, other official specialties expect pediatrics, and the number of beds.[24] As municipality-level covariates, we control for demographic variables of the municipalities such as population and the share of children aged under 15 years old.

4.4 Sample Construction

In constructing the sample, the polled number of clinics in 5 waves of the SMI is 490,127. Among them, 122,490 clinics are identified as "pediatric clinics" whose official specialties included "pediatrics". After excluding missing values, the potential sample consists of 113,470 clinics. Because this paper uses the clinics that are merged with the MSCI eligible age, the number of clinic-year observations is 62,221 in the main specification. A flowchart of the sample construction is summarized in Figure 2.

5 Econometric Specification

5.1 Difference-in-Differences-in-Differences

To identify the effects of the MSCI expansion, we first consider the difference-in-difference (DD) framework that exploits regional differences in the MSCI eligible age (i.e., first difference) and their changes over years (i.e., second difference). However, we should carefully address potential endogeneity in the MSCI expansion. While there is little literature that examines the determinants of MSCI expansion, numerous regional factors may be associated with it. For example, the fiscal condition of the municipalities (Ando, 2017) and timing of mayoral elections (Adachi & Saito, 2016) seem to affect the MSCI eligible age.

Because these municipality level characteristics may also affect the behavior of pediatricians even in the absence of MSCI expansion, we consider that the naive DD may not fully address municipality level unobservables that jointly affect MSCI eligible age and pediatrician behavior. Thus, we took advantage of the differential effects of the MSCI according to the types of clinics, in the same spirit as Garthwaite (2012).

Our triple-difference strategy is based on the following stylized facts in Japan. In Japan, physicians who learned pediatrics as their primary specialization tend to choose only "pediatrics" as their official specialty, without choosing "internal medicine", which indicates primary care for adult patients. Because

[24] In Japan, medical institutions with less than 19 beds are classified as a clinic and those with more than 20 are classified as a hospital.

the main customer base of these clinics is necessarily children, they would have been greatly affected by the MSCI expansion. In this paper, we call them "child clinics". On the other hand, there are other clinics that provide primary care for adults as well as children. In Japan, these clinics generally practice both "internal medicine" and "pediatrics". Because children are only a portion of customers for these clinics, we can reasonably assume that the MSCI expansion has increased the total number of patients in child clinics more than in all-generation clinics, even though these clinics share the similarity in many aspects.

Based on the discussion above, our preferred specification includes all pediatric clinics that advertise "pediatrics" on their signs, excluding other clinics. The following equation is then estimated as a baseline model,

$$y_{it} = \alpha_0 Child_{it} + \alpha_1 Elig_{mt} + \alpha_2 * Child_{it} * Elig_{mt} + \alpha_3 X_{it} + \alpha_4 Z_{mt} + \theta_m + Year_t + \psi_{it}, \tag{22}$$

where y_{it} is outcome variables in clinic i in year t; $Child_{it}$ is a binary variable for child clinics; $Elig_{mt}$ is the MSCI eligible age of the outpatient care in municipality m in year t. If clinic i was located in municipality m in year t, $Elig_{mt}$ is allotted to this clinic. X_{it} is a vector of clinic level covariates; Z_{mt} is a vector of municipality level time varying covariates; θ_m is municipality fixed effects; $Year_t$ is year effects. Note that municipality level time varying covariates are controlled for, even though they are supposed to be endogenous to the MSCI expansion. For example, inter-municipality migration (i.e., household with children can move to the municipality that offers a generous MSCI) may indicate that population, which is the most basic characteristic of a municipality, is also endogenous.[25] However, we assume that endogeneity of municipality level variables is limited because MSCI is only relevant for healthcare utilization. With X_{it}, we assume clinic ownership and existence of beds are mostly a pre-determined covariate for the MSCI expansion because clinics do not change their owners according to the generosity of the MSCI.[26]

In this equation, α_1 captures municipality level heterogeneous trends that are associated with the MSCI expansion. These factors may lead to spurious results on the effects of the MSCI if they are not correctly controlled for. However, we assume unobservable municipality level factors may affect child clinics and all-generation clinics together. If this is the case, the interaction term ($Child_{it} * Elig_{mt}$) captures the differential effects on child clinics from the MSCI expansion, which can be regarded as the causal effects from the MSCI. In short, our identifying assumption is that the MSCI expansion was uncorrelated to unobservable factors that could have generated clinic-level variation in outcome variables, conditional on time invariant municipality- and clinic-level characteristics and prefecture-specific trends.

[25] Nakajima & Tanaka (2014) examine how local-government-sponsored pro-natal policies, which include the MSCI, affect fertility in Japan and find that self-selected migration across municipalities by parents may cause substantial upward bias on the estimated impacts on fertility. This suggests that parents who want more children may migrate to the municipalities that offer generous pro-natal public services.

[26] Some clinics with fewer than 20 beds are classified as clinics in Japan. The main purpose of having such a small number of beds is to provide for relatively short stays for emergency patients with mild conditions. Because these beds are used for inpatients, the extension of MSCI eligible age for outpatient care, which is the treatment variable in this study, has no direct influence.

5.2 Potential Threats for Identification

5.3 Endogeneity of MSCI expansion

Our triple differences analysis makes a common trend assumption more plausible than the naive DD, but there may be two problems with our identification. First, unobservable municipality level characteristics may affect the generosity of MSCI and, at the same time, they may have differential effects on child and all-generation clinics. For example, it is possible that parents with higher incomes are more likely to visit child clinics because these clinics provide specialized care for children, and the income level of the municipality is also associated with the expansion of MSCI. In this case, we cannot assume that child clinics and all-generation clinics follow similar trends in the absence of MSCI expansion.

To address this, we explicitly consider the determinants of the MSCI expansion from 1999 to 2011. First, we allow the independent effects from municipal level and prefecture level characteristics because the MSCI is implemented as a joint subsidization program of municipal and prefectural government. Second, we focus on the within-prefecture variation of the MSCI eligible age after controlling for prefecture-level fixed effects. Finally, we begin exploring the determinants of MSCI eligible age by estimating the following equation,

$$Elig_{m,2011} - Elig_{m,1999} = \upsilon_0 + \Upsilon_1 Z_{m,1999} + \Upsilon_2 Pref_p + \epsilon, \tag{23}$$

where $Elig_{m,2011} - Elig_{m,1999}$ is the difference of the MSCI eligible age between 2011 and 1999, υ_0 is a constant term, $Z'_{m,1999}$ is a vector of municipality-level characteristics as of 1999, and $Pref_p$ is a vector of a binary indicator for prefectures, ϵ is an independent and distributed error term. With $Z_{m,1999}$, we control for several basic characteristics of the municipalities such as demographics, income level and industrial structure. Among these variables, we find that the shares of nuclear households and of one-person households are negatively associated with the MSCI expansion from 1999 to 2011, while income, land price, the number of pediatricians and industrial structures may also affect for the MSCI expansion with less precise effects.[27]

Whatever the reasons, the endogeneity of the MSCI expansion is relevant for our identification. Therefore, we control for the direct effects from the regional characteristics in the baseline year by additionally controlling for prefecture specific linear trends and the interaction terms of time trend and municipality-level characteristics as of 1999, accordingly:

$$\begin{aligned} y_{it} = \alpha_0 Child_{it} + \alpha_1 Elig_{mt} + \alpha_2 * Child_{it} * Elig_{mt} + \alpha_3 X_{it} + \alpha_4 Z_{mt} + \theta_m + Year_t + \\ \alpha_5 T * Pref_p + \alpha_6 T * Z_{m,1999} + \psi_{it}, \end{aligned} \tag{24}$$

where, $T * Z_{m,1999}$ absorbs heterogeneous trends by the municipality-level characteristics as of 1999 and $T * Pref_p$ absorbs prefecture specific trends. In this main analysis, we assume the MSCI expansion

[27]For further results, see Online Appendix B.

from 1999 to 2011 was sufficiently exogenous, conditional on these trend terms and other predetermined covariates.

5.4 Endogeneity of Specialty Choice

Even if the MSCI eligible age can be assumed to be conditionally independent, we have another threat for the identification. Second threat is the endogeneity of $Child_{it}$, which is a binary variable for child clinics. This endogeneity comes from the fact that all clinics in Japan freely choose their specialties. Therefore, it is possible that physicians who have no experience with child patients advertise "pediatrics" on their signs, while the choice of official specialties is generally determined by education and training in the medical university. Regardless of the actual reasons, the expansion of the MSCI ($Elig_{mt}$) can affect the probability of choosing to become a pediatric clinic, namely $Child_{it}$, which generates bias on the coefficient of the triple differences term (i.e., $Child_{it} * Elig_{mt}$). In the Online Appendix C, we carefully explore the association between $Child_{it}$ and $Elig_{mt}$ directly, but find no statistically significant associations. The specialty of physicians is likely mainly determined by education during medical university and not responsive to cost-sharing policy.

6 Results

6.1 Descriptive Statistics

Table 1 provide descriptive statistics for outcome variables in the full sample and sub-samples. Columns (1) and (2) report the results from all pediatric clinics in the SMI. The total number of clinic-year observations is 90703, but we cannot implement statistical analysis with this full sample because the clinics that are not matched with our original survey on the municipality-level MSCI eligibility must be dropped. Even though we have to drop many pediatric clinics, the final sample includes 62,221 clinics because most large municipalities are covered.[28] In columns (3) and (4), the descriptive statistics are reported. In comparison with the results of the whole sample reported in columns (1) and (2), the characteristics of the main sample do not seem to change greatly: The number of patients per month is about 1000 in both samples and the number of consultation days is also similar.[29] In columns (5) and (6), we present the means and standard deviations in the sample of child clinics that are supposedly affected by the MSCI expansion. In addition, columns (7) and (8) report the results of all-generation clinics that serve both adults and children. In comparing these two groups, we find that the number of first visits is larger in child clinics while they have generally similar characteristics.

Table 2 provides descriptive statistics for covariates in the sample of child clinics in columns (1) and (2) and all-generation clinics in columns (3) and (4). The two groups exhibit little difference in the ownership structure of the clinics, but all-generation clinics declare more clinical specialties than child clinics. For example, 31% of all-generation clinics practice "gastroenterology" because diseases of the digestive system

[28] 25% of the clinics are dropped.

[29] The statistical test of mean differences found *statistically* significant differences in most of the variables, but the differences still appeared to be small.

are prevalent in Japanese adults. Also, the proportion of child clinics with the "Allergy" specialty is 20 %, which higher than that of all-generation clinics. On the municipality-level covariates, child clinics seem to be located in large municipalities in terms of population.

6.2 Common Trends

Before introducing the regression results, we checked the trends of the main outcome variables in Figure 4. In this figure, we plotted the average values of the number of visits, population within a 3 km radius, the number of consultation days by survey year, types of clinics (i.e., child vs all-generation) and the extent of MSCI expansion. As noted in Section 2, some municipalities restricted the MSCI eligible age to preschool children, but others expanded it much further. We therefore categorized the municipalities that extended MSCI to cover school age children in 2008 as "rapid extenders" and those who extended MSCI only for preschool children at that time as "slow extenders". In Figure (a), the number of visits at child clinics in the "rapid" municipalities followed a similar trend as child clinics in "slow" municipalities from 1999 to 2002, but the former exceeded the latter from 2005 to 2011 when the MSCI was rapidly expanded. In addition, we find the trends of all-generation clinics did not deviate from those of child clinics, but there is a sharp increase in only the child clinics in "rapid" municipalities.

In Figure (b), we found suggestive graphical evidence of the MSCI effects on location choice. In this figure, we show the subsample results focus on the newly established clinics because existing clinics do not move due to high fixed costs. Also, we show the medians rather than means in this figure because means are fluctuated by the data of some clinics which are located in extremely dense area. The trends of the 4 groups were very similar from 2002 to 2005, but population levels in the location of all generation clinics decreased in 2008. Child clinics in the "slow" municipalities also saw slightly decreased population levels in their practice location. However, child clinics in the "rapid" municipalities, which were exposed by large extension of MSCI eligible age, did not. This suggests that MSCI incentivized pediatric clinics to locate to more densely populated areas. In figure (c), we examined trends in the number of consultation days. Although we did not find clear results for these outcomes, it should be noted that the number of consultation days was the lowest in the child clinics in "rapid" municipalities in 2011.

6.3 Effects on the Number of Visits

Regression results on the number of visits, which is the total number in the clinic in October in the survey years, are presented in Table 3. In column (1), we report the results with the minimum number of covariates (i.e., year effects and municipality fixed effects). In this specification, the coefficient $Elig_{mt}$ absorbs differential trends in the number of visits that are associated with the MSCI expansion. Because unobservable time trends across municipalities during the study periods may contaminate the MSCI effects with unobserved time-varying factors, the coefficient of $Elig_{mt}$ may provide biased estimates of the effects of the MSCI expansion. However, after controlling for $Elig_{mt}$, the interaction term ($Child_{it} * Elig_{mt}$) uncovers the effects of the MSCI on child clinics. The coefficient of the interaction term is 6.219 ($p < 0.01$), indicating that the MSCI increases the number of patients treated by clinics.

The estimated effects are further refined by adding covariates. In column (2), results with full covariates are presented in Table 2. The coefficient of the interaction term is 6.207 and statistically significant. The point estimate indicates that extension of the MSCI eligible age by 10 years old increases the number of visits to pediatric clinics by 62. If the eligible age is extended by 10 years, because the MSCI eligible age was raised by about 10 years from 1999 to 2011, the number of visits to pediatric clinics will increase by 5.9% (93 people per month). This suggests that the price elasticity of health care demand among children is about 0.059,[30], which is lower than the estimate from the RAND health insurance experiment (Manning *et al.*, 1987) but is compatible with previous studies from patient level data in Japan (Shigeoka, 2014; Fukushima *et al.*, 2016; Takaku, 2017; Iizuka & Shigeoka, 2018).[31] In particular, Iizuka & Shigeoka (2018) reveal that the price elasticity of healthcare demand among children is approximately half that of adults. Our estimate using clinic-level data is generally consistent with the most credible study on the impacts of the MSCI. In addition, the results are fairly robust when we control for clinic fixed effects in column (3), instead of municipality fixed effects in column (2). Again, the coefficient of the interaction term is very similar to that in column (2) and statistically significant.

In Table 4, we break down the total effects into first visits, follow-up visits, and off-hour visits. Columns (1) to (3) report the results with the municipality fixed effects, but columns (4) to (6) report those with clinic fixed effects. In both specifications, the main findings are very similar. First, we find clear significant effects on the number of first visits in columns (1) and (3), indicating that the MSCI increases the health care demand of patients with acute conditions. Second, the MSCI may increase the number of follow-up visits but the coefficient of the interaction term is not so precise. Finally, for the number of off-hour visits, we do not find any increases in columns (3) and (6).

6.4 Effects on Practice Location Choice

Next, we show how the MSCI expansion changes physician practice location choice. In constructing a variety of measures of location characteristics, we first report the results on the population density within a radius of 3 km from the location of the clinic as a benchmark case. Results are reported in Table 5. In this table, columns (1) to (3) show the results from the full sample and columns (4) and (5) report the results of the subsample analysis that focuses on newly opened clinics. Because we identify new clinics by checking their existence in the previous survey, the analysis in columns (4) and (5) uses the data in 2002, 2005, 2008, and 2011. In column (1), the result of the minimum covariates is reported. The coefficient of the interaction term is negative and not statistically significant. The result is the same when we control for full covariates in column (2). The reason why we find non-significant effects may be due to high fixed costs. Given that clinics rarely move once they are opened due to high moving costs (Polsky *et al.*, 2000), we cannot expect that operating clinics move into municipalities with generous MSCI.

[30] Because the MSCI generally provide free care, out-of-pocket expense decreases 100%. Therefore, price elasticity is -0.059 (= 5.9%/ -100%).

[31] Note that half of the municipalities that implement MSCI impose a slight copayment (Ministry of Health & Welfare, 2013), thereby the reduction of the denominator (i.e., 100%) is somewhat exaggerated. Price elasticity would therefore be much higher if the copayment reduction is measured with complete accuracy.

However, when the sample is restricted in the newly established clinics, the effects of the MSCI seem to be sizable: In column (5), the coefficients of the interaction term are 70.435 and statistically significant at a 99% confidence interval. This suggests that an increase of the eligible age by 10 years increases the population within a radius of 3 km from the location of the clinic by 704.35 persons, which is 10% of the mean number of child clinics.

These results are consistent with the theoretical prediction of Section 3. In fact, contrary to the results of other studies (Yang *et al.*, 2013; Chen *et al.*, 2017; Huh, 2017), newly established clinics choose more densely populated areas as their destination of practice when a medical subsidy program is expanded. Although we find no evidence of the migration of operating clinics, the generosity of health insurance has fairly sizable effects on the location decision of new clinics. This indicates that health insurance does not immediately change the geographical distribution of clinics, but we should not ignore the long-run consequences on the geographical distribution of physicians.

We also check the robustness of the results by using an alternative measure of the location's urbanity. Table 6 summarizes the results. From columns (1) and (3), we show the results with and without covariates. In addition, the definition of each measure is shown in the leftmost column. Here, the results remain unchanged if we adopt the population density within 1 km or 5 km. In addition, the coefficient of the interaction term is positive and statistically significant for the population density of the SAA.

Next, we construct a binary variable of the location. This variable takes a value of one if a clinic is located in "rural" area,[32] as is key dependent variables in the previous studies (Chen *et al.*, 2017; Huh, 2017). When we investigate the impact of location choice with this crude measure, any statistically significant effects are not found.

Also, we can evaluate the effects by age groups of the population. This analysis is of particular importance when we consider the underlying mechanisms on why physicians choose more densely populated areas. One potential reason is physicians' preference on city amenity as is postulated in our model, but it is also likely that there are more young couples with children in these areas and higher demand for child health care due to the subsidy expansion attracts more physicians. Though this likelihood would be almost denied by the interpretation in the theoretical model Section 3 showed, we can test these different hypothese in this supplemental analysis. The results, shown in lower raws in Table 6, indicate that physicians do not respond to the child population at all. In all empirical models, coefficients of the interaction term are not statistically significant when the child population is used as an outcome variable. Instead, they choose to operate in the areas with higher number of adult population, suggesting that physicians are attracted in urban areas not because of higher demand but because of other potential merits which are attached to these areas (i.e. city amenity). This interpretation is identical to one in the theoretical model, and one intuition to support this interpretation is that many children live in housing complex in Japan, but pediatricians do not always open their clinics in the first floor of these large buildings because they are generally constructed in a bit remote areas from downtown. Rather, they may choose to operate in much more bustling areas even though it is not convenient for sick children and their parents to visit.

[32] If the population density of a 3 kilometer radius is less than 25 percentile, the clinics is categorized in "rural" group.

6.5 Effects on the Number of Consultation Days

Although MSCI expansion increases the number of visits per clinic and induces new physicians to open their clinics in urban areas, they can also change working hours. Previous studies have examined this effect. For example, Enterline *et al.* (1973) and Garthwaite (2012) found that physicians may reduce working hours when health insurance is expanded. In particular, Enterline *et al.* (1973) emphasized that physicians could earn as much or more after Medicare was introduced by working shorter weeks. This interpretation is consistent with the argument that physicians' labor supply curve is backward bending (Feldstein, 1970; Rizzo & Blumenthal, 1994; Thornton & Eakin, 1997)[33]

Our theoretical model ignores the choice of working hours for the simplification, but actual physicians jointly decide "where" and "how long" they work. Though this joint determination may be too complex to be incorporated in our theoretical model, we can consider some consequences. The MSCI expansion increases the demand for health care services in both rural and urban areas, which reduces working hours, as previous studies show. On the other hand, the MSCI expansion induces physicians to locate their clinics in urban areas, which would weaken profit increases in urban areas and the effect of reduced working hours.

The results are reported in Table 7. In this table, Panel A reports the baseline results with a full sample including municipality-level fixed effects. Clinic-level fixed effects are controlled for in Panel B. For analysis of practice location choice, statistically significant effects are not observed in Panel B. Panel C, reports the results from the subsample analysis that focus on new clinics.

In addition, we create binary variables that examine whether physicians open their clinics more than 4, 5, and 6 days in columns 2 to 4. In column 1 of Panel A, the coefficient of the interaction term is negative and statistically significant, suggesting that physicians reduce the number of consultation days when MSCI is expanded. Also, the probability of opening clinics more than 4 days per week decreases in column (2). Although these results suggest that physicians in child clinics reduce the total number of consultation days, the quantitative impacts are small. For example, the coefficient of the interaction term is only -0.008, indicating that the expansion of the MSCI eligible age from 0 to 15 years old reduces the number of consultation days by 0.02 days. In addition, these results are not so robust when we control for clinic level fixed effects in Panel B. Even among the subsample of newly established clinics, we do not find statistically significant effects from MSCI expansion. We do not find a robust reduction of the number of consultation days, while signs of the effect are consistent with the experience in Quebec (Enterline *et al.*, 1973) and the US (Garthwaite, 2012). In Online Appendix D, we present more detailed results. In short, we find a slight reduction in opening hours and the probability of operating for each day in a usual week, but the quantitative impacts appear negligible.

[33] Even though these two papers found suggestive evidence on the backward bending labor supply curve among physicians, a literature review by Nicholson & Propper (2011) concludes that "there appears to be general agreement that as with most occupations, physicians are not particularly responsive to wage changes. And income elasticities are small."

7 Placebo Tests

In the same spirit as Abadie *et al.* (2010), we check the robustness of our main results by implementing placebo tests. Our placebo tests are implemented using the following procedure. First, we randomly replace MSCI eligible age only for child clinics (i.e., treatment group), while leaving the MSCI eligible age of all-generation clinics (i.e., control group) unchanged. In this step, we build a sample *true* control group and treatment group which are assigned a *placebo* MSCI eligible age that was randomly chosen from the actual distribution of the MSCI eligible age. Second, we implement the same triple differences analysis for that sample. If our main results are obtained just by chance, this placebo test also provides statistically significant effects of the MSCI expansion. Next, we implement steps 1 and 2 300 times, and store t statistics of the interaction term coefficient. Finally, we compare true t statistics with the placebo t statistics histogram.

Figure 5 presents the results for our main outcomes, namely the number of visits, the number of consultation days and population within a radius of 3 km. In all outcome variables, our placebo tests seldom replicate the true results. For example, true t statistics in Figure 5-(a) exceeds 3, which indicates strong positive impacts. However, most of the placebo t statistics are concentrated around 1.5. Note that some placebo t statistics may indicate statistical significance because the *placebo* MSCI eligible age for the treatment group often takes a value very close to the true one, just by chance. This is because the MSCI eligible age is a discrete variable that ranges from 0 to 18, as is shown in Table A2. However, while some placebo tests provide statistically significant effects, we could not frequently replicate the *true* results with the randomly assigned MSCI eligible age. Therefore, we confirm that our main findings are unlikely to be driven by chance.

8 Conclusion

Although many developed countries that have already achieved universal health coverage follow a general trend to reduce the share of out-of-pocket expense (Fan & Savedoff, 2014; Dieleman *et al.*, 2017) and there is growing interest in supply-side responses to health insurance expansion (Finkelstein, 2007; Kondo & Shigeoka, 2013), no study has investigated how primary care physicians respond to large-scale reduction in patient cost-sharing. In addition, by theoretical consideration, we show that physicians' responses in practice location choice greatly differ according to the margins of health insurance expansion. In accordance with previous findings (Yang *et al.*, 2013; Chen *et al.*, 2017; Huh, 2017), physicians are more likely to choose rural areas as their practice location when health insurance has been expanded in the extensive margin. However, standard economic theory predicts that expanded generosity of health insurance may induce physicians to concentrate in densely populated areas, which worsens a common policy issue in developed countries; that is, the geographical mal-distribution of physicians (Ono *et al.*, 2014).

By exploiting recent dramatic expansion of a subsidization program for children's health care utilization, the MSCI, we examine how primary care physicians respond to changes in demand for side incentives. Data used in this paper are more comprehensive than those used in recent previous studies (Garthwaite, 2012;

Buchmueller *et al.*, 2016) and the institutional setting in Japan may make the interpretation of the results more straight-forward than the results from the US, where providers are subject to various health insurance plans with different reimbursement rates. In addition, by using geographical information on the location of clinics, we construct more accurate measures of location characteristics than previous papers (Escarce *et al.*, 1998; Polsky *et al.*, 2000; Yang *et al.*, 2013; Chen *et al.*, 2017). To reduce potential bias from observational data, difference-in-differences-in-differences analysis are employed in a similar spirit to Garthwaite (2012).

Our analysis suggests that the MSCI expansion increases the number of monthly visits treated by clinics. Although we show the results with municipality fixed effects as preferred specification, this result remains unchanged when controlling for clinics' fixed effects. For quantitative impacts, full implementation of the MSCI increases the number of visits by 5.9%. Consistent with recent rigorous study on the impact of MSCI (Iizuka & Shigeoka, 2018), this paper indicates that the price elasticity of primary care utilization among children is slightly lower than that reported in RAND HIE (Manning *et al.*, 1987).

We also investigate the effects on location choice and the number of consultation days. Importantly, we find suggestive evidence that generous health insurance system accelerates the concentration of physicians into densely populated area. The results show that the population density where a newly established clinic is located, which is considered as a proxy for the urbanity of the location, increases by 18% when the MSCI is expanded to cover all children (i.e., up to 15 years old). Consistent with Escarce *et al.* (1998) and Polsky *et al.* (2000), we find the location choice of operating clinics are not responsive to the MSCI expansion but newly established clinics may change the practice location sensitively. These results are also consistent with our core-periphery model on physician's practice location choice. In addition, we should note that our results on this outcome are in sharp contrast with the findings from previous studies exploring the impacts of the expansion of health insurance in the *extensive margin*. For example, Yang *et al.* (2013), Chen *et al.* (2017) and Huh (2017) consistently found that the health insurance expansion for the uninsured population may alleviate geographic mal-distribution of physicians since healthcare costs, which are the sources of physicians' income, increase mostly in deprived areas where many previously uninsured persons live. However, this study deals with the expansion of the *intensive margin* of public health insurance. In the MSCI expansion, all insured children experienced a large reduction in coinsurance if their age is lower than the threshold age. This enhances income levels of physicians even in the competitive urban areas, which induces physicians who prefer city life to choose to open their clinics in such urban areas, accelerating the geographical mal-distribution of primary care physicians. To support this interpretation, we build up a formal theoretical model to compare the effects of health insurance on physicians' practice location choice between intensive and extensive margins.

Of course, it is possible that MSCI inflates children's health care demand in densely populated areas and this attracts more pediatricians. However, we find no evidence which supports this story. Rather, pediatricians open their clinics in the areas where many adults, not children, lives. We interpret these results as pediatricians choose practice location according to their taste on city amenity, irrespective to health care needs. Taken together, free health care for children is far from an effective policy to provide access to health services in deprived areas because these areas become less likely to be chosen as the physicians' practice location. Also, policy makers in countries with universal health coverage should be aware of the possibility that expansion of the health insurance in the *intensive margin* may lead to opposite consequences with the expansion in the *extensive margin*.

Additionally, we found suggestive evidence that physicians slightly reduce their number of consultation days per week when the MSCI is expanded. This result is consistent with Enterline *et al.* (1973) and Garthwaite (2012); however, the quantitative impact is small and almost negligible. Given that the total monthly number of visits exhibits sizable increases, the slight reduction in consultation days may indicate that patients receive shorter consultations under a generous MSCI, which results in lower quality of care, but this point should be carefully examined using other datasets that contain precise data on working hours.

This paper sheds new light on physicians' responses to health insurance expansion by using accurate data on clinic location, but there are some limitations. First, due to incomplete responses to our original survey on the MSCI eligibility criteria (Takaku, 2016), some municipalities are excluded from the empirical analysis. In terms of the number of clinics, approximately 37% of the clinics are excluded. Although this may limit the external validity of our analysis, the characteristics of the excluded clinics seem to be sufficiently similar to those used in the analysis, which suggests that the potential problem from this limitation is not serious. Second, we focus on the behavior of clinics rather than physicians in this paper. Because multiple physicians work for a clinic, we should further explore responses of individual physicians. For example, it is possible that responses to the health insurance expansion differ according to physician gender. Uncovering detailed heterogeneous effects posits a great avenue for future studies.

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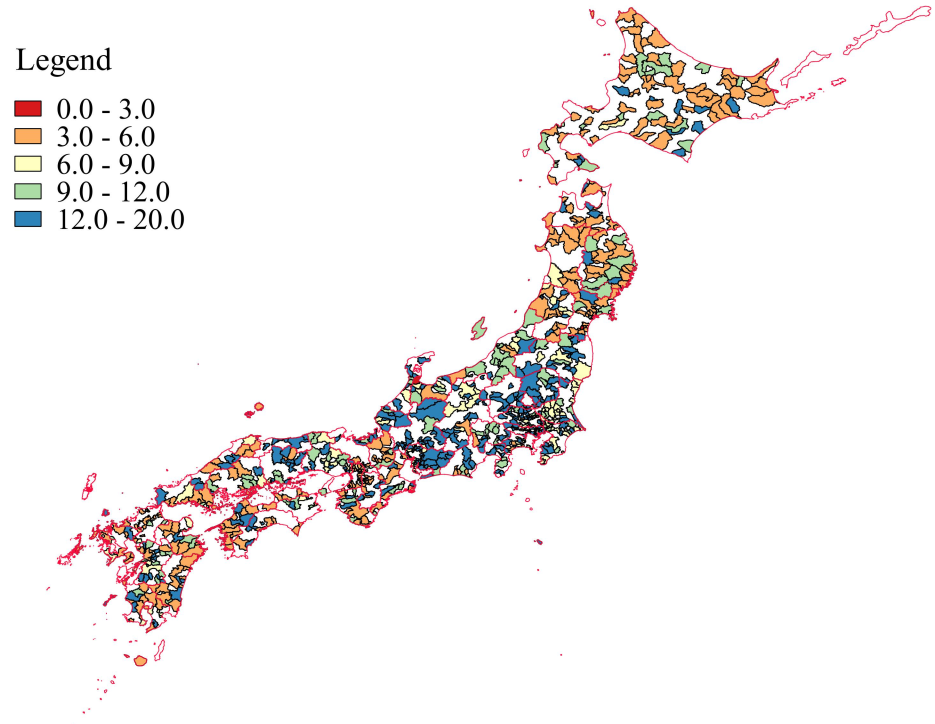


Note: The figure represents the map as of 2011. Municipalities shown in white are those that did not answer the original survey on the MSCI's eligibility criteria. Other colors represent MSCI eligible age. For example, municipalities with an eligible age of 3–6 years old are shaded in orange. Those with an eligible age of 12–20 years old are shaded in blue. Okinawa prefecture is excluded due to space limitations.

Figure 1: Municipality Level Map of MSCI Eligible Age: 2011

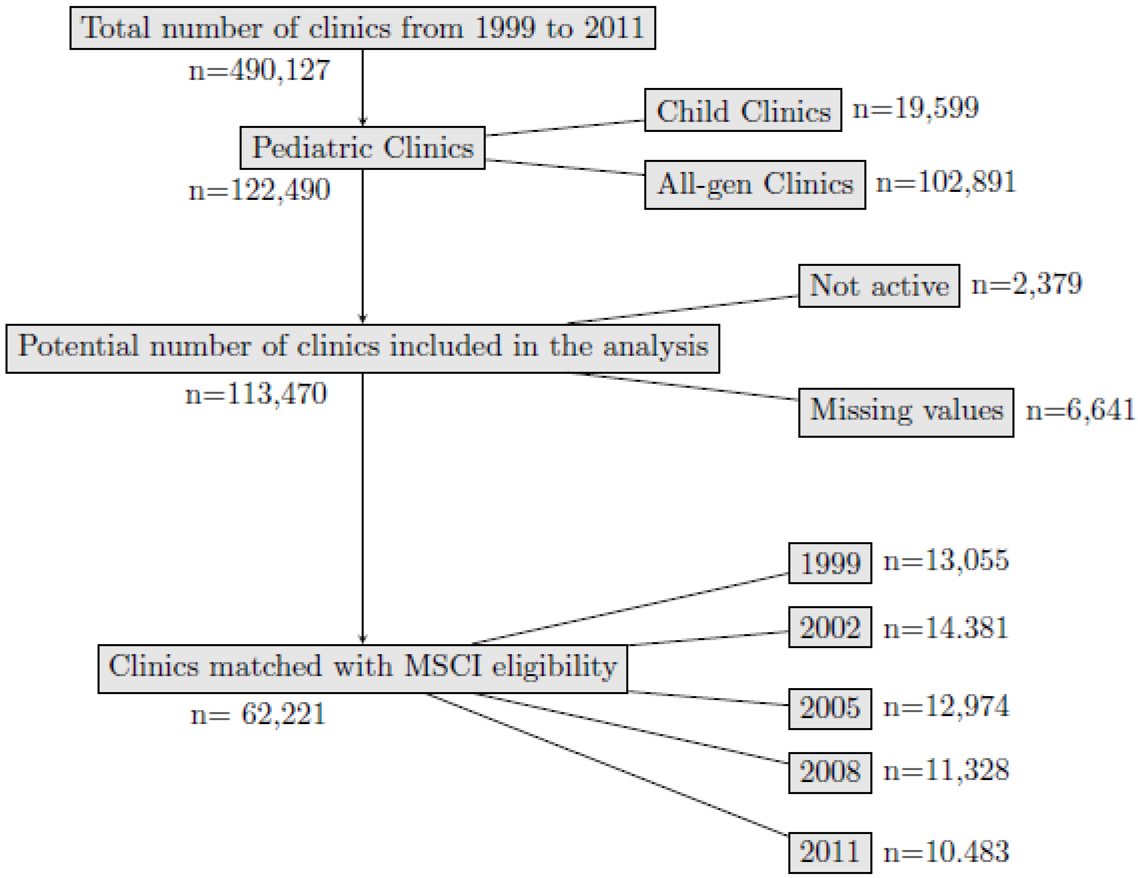


Figure 2: Flowchart of the Sample Construction

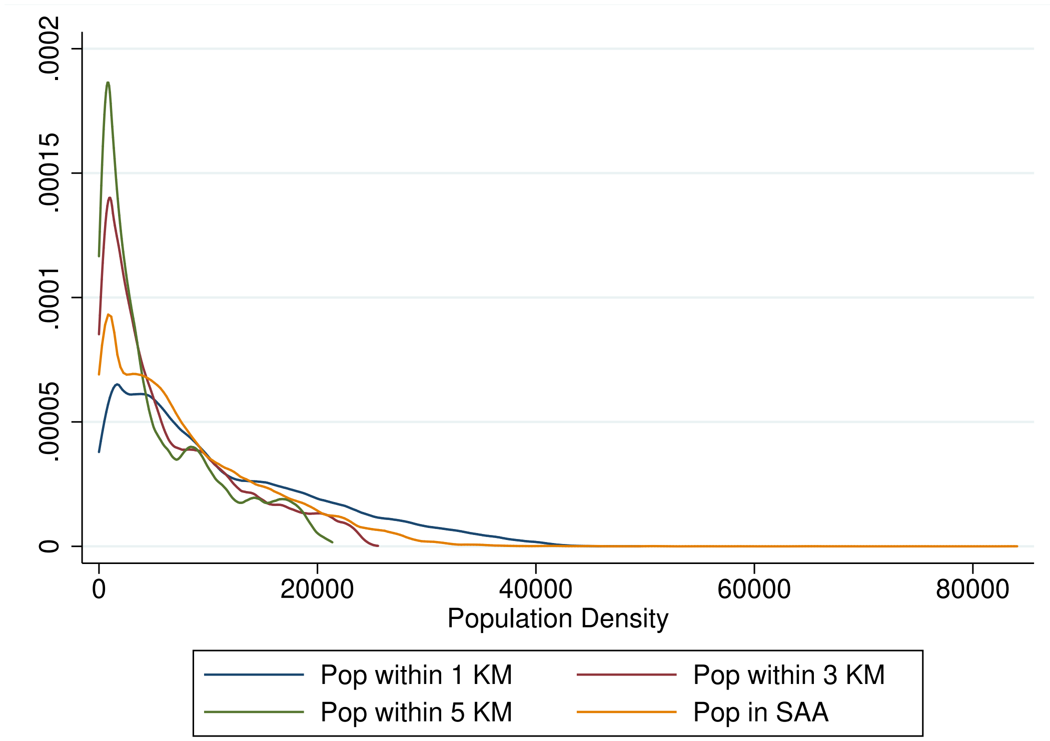
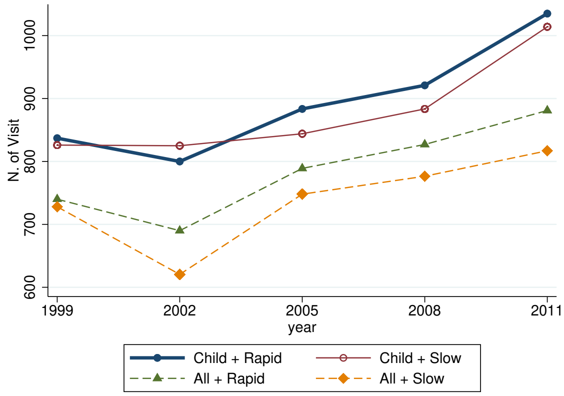
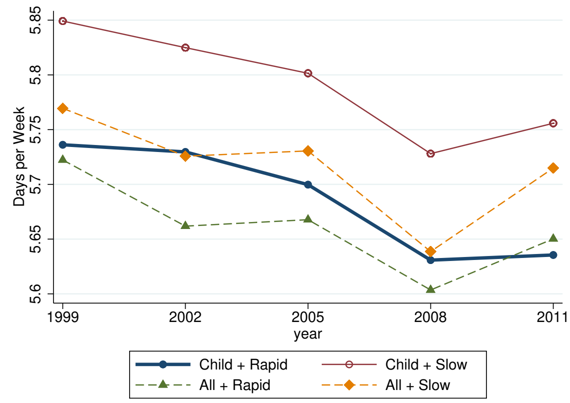


Figure 3: Distribution of Population Density Around the Clinic's Location

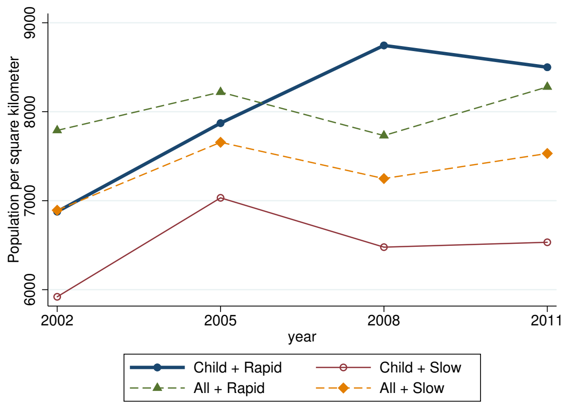
Note: Population density is calculated as population per km^2 where each clinic is located. "SSA" represents the smallest administrative area (in Japanese, *choume*).



(a) Number of Visits



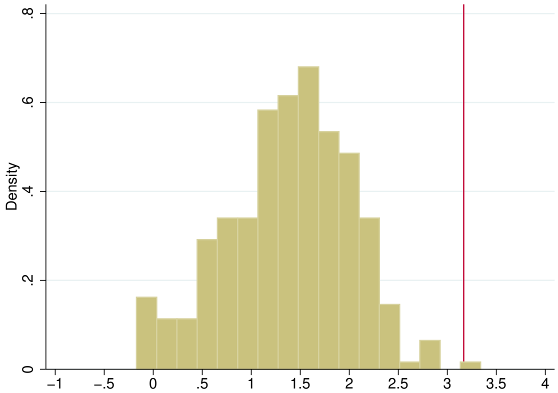
(c) Number of Consultation Day



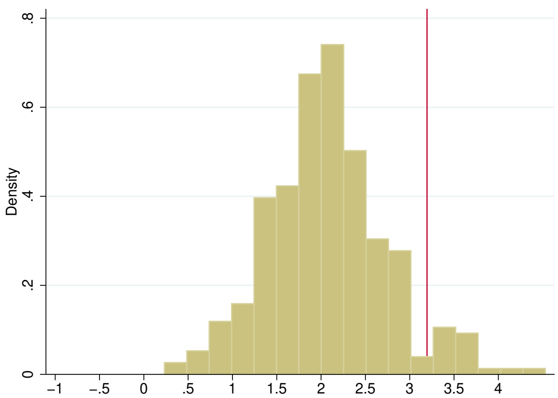
(b) Population within 3 KM

Note: “Child” represents child clinics and “All” represents all-generation clinics. Clinics are classified as “Rapid” if they are located in the municipalities that extended the MSCI eligible age for school age children in 2008, and “Slow” otherwise.

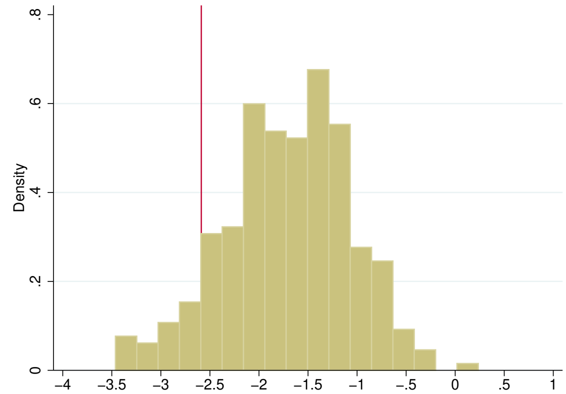
Figure 4: Trend of Main Outcome Variables



(a) Number of Visits



(b) Population Density : 3 KM Radius



(c) Number of Consultation Day

Note: Vertical line represents t statistics of the interaction term coefficient ($Elig{*}Child$) in the main specifications with *true* sample. Histogram represents the distribution of the t statistics from 300 placebo tests.

Figure 5: Placebo Tests

Table 1: Descriptive Statistics: Main Outcome Variables

	All		Sample Included in Analysis		Child		All-Gen	
	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Number of Monthly Visits								
Total Visits	1014.81	[1,057.62]	982.86	[1,002.37]	1036.01	[831.83]	970.84	[1,036.70]
First Visits	125.63	[192.46]	123.06	[188.76]	196.95	[251.92]	106.35	[166.79]
Follow-up Visits	889.19	[986.00]	859.79	[926.54]	839.06	[771.59]	864.49	[958.07]
Off-hour Visits	17.13	[76.20]	17.17	[74.04]	27.17	[99.17]	14.75	[66.33]
Practice Location Choice								
Population within 1 KM	11135.74	[9,281.10]	13278.12	[9,233.43]	12654.98	[7,828.74]	13419.10	[9,516.92]
Population within 3 KM	6434.23	[6,005.08]	7774.68	[5,975.81]	7095.78	[5,003.74]	7928.26	[6,164.21]
Population within 5 KM	5294.58	[5,252.17]	6434.03	[5,250.12]	5742.96	[4,462.23]	6590.37	[5,400.27]
Population Density in SAA : Total	8246.94	[7,274.97]	9849.33	[7,424.24]	9489.67	[6,960.08]	9930.69	[7,522.96]
Population Density in SAA : age $<$ 15	956.64	[870.97]	1140.30	[905.79]	1168.03	[931.58]	1134.03	[899.75]
Population Density in SAA : 15 $<=$ age $<$ 65	6333.66	[5,817.48]	7568.73	[5,948.90]	7153.62	[5,477.24]	7662.64	[6,046.60]
Population Density in SAA : age $>$ 65	1771.89	[1,572.33]	2091.62	[1,634.98]	1877.06	[1,417.57]	2140.17	[1,676.48]
Rural	0.25	[0.43]	0.14	[0.34]	0.08	[0.27]	0.15	[0.36]
Consultation Day								
Total Consultation Days	5.59	[1.06]	5.57	[1.07]	5.65	[0.89]	5.56	[1.11]
Consultation Days $>=$ 4	0.95	[0.22]	0.95	[0.22]	0.97	[0.17]	0.95	[0.22]
Consultation Days $>=$ 5	0.74	[0.44]	0.73	[0.45]	0.74	[0.44]	0.72	[0.45]
Consultation Days $>=$ 6	0.02	[0.12]	0.02	[0.12]	0.01	[0.12]	0.02	[0.12]
Obs.	113470		62,221		11,479		50,742	

Note: In the empirical analysis, clinics that are not matched with the authors' original survey on the MSCI system are dropped. Number of outpatients is the total number in October in each survey year. Population density is calculated as population per km^2 where each clinic is located. Number of consultation days is per week, ranging from 0 to 7 days. "Child" represents child clinics that mainly serve children. "All-Gen" represents all-generation clinics that serve adults as well as children.

Table 2: Descriptive Statistics: Covariates

	Child		All-Gen			
	Mean	S.D.	Mean	S.D.	Diff	
Clinic Level Covariates						
Government-Owned Clinics	0.01	[0.12]	0.03	[0.16]	-0.01	***
Other Public Clinics	0.00	[0.02]	0.00	[0.03]	0.00	
Cooperative Clinics	0.39	[0.49]	0.38	[0.48]	0.01	***
Clinics Owned by Solo-practice Physician	0.60	[0.49]	0.60	[0.49]	0.00	
Beds	0.08	[0.27]	0.11	[0.31]	-0.03	***
Official Specialty						
Respiratory Medicine	0.01	[0.08]	0.14	[0.35]	-0.14	***
Gastroenterology	0.01	[0.12]	0.31	[0.46]	-0.29	***
Cardiology	0.01	[0.12]	0.23	[0.42]	-0.22	***
Neurology	0.01	[0.08]	0.03	[0.17]	-0.03	***
Allergy	0.20	[0.40]	0.09	[0.28]	0.12	***
General Surgery	0.02	[0.12]	0.13	[0.34]	-0.12	***
Orthopedic Surgery	0.02	[0.14]	0.07	[0.25]	-0.04	***
Obstetrics and Gynecology	0.06	[0.24]	0.03	[0.18]	0.03	***
Gynecology	0.02	[0.13]	0.03	[0.16]	-0.01	***
Ophthalmology	0.01	[0.09]	0.02	[0.13]	-0.01	***
Dermatology	0.04	[0.19]	0.15	[0.36]	-0.11	***
Radiology	0.00	[0.05]	0.11	[0.31]	-0.11	***
Municipality-Level Time Varying Covariates						
Total Population	645,362.00	[856923]	557,874.00	[791511]	87,488	***
Proportion of Children Aged Under 15 Years	0.14	[0.03]	0.14	[0.04]	0.00	***
Proportion of Elderly Aged Over 65 Years	0.19	[0.02]	0.19	[0.02]	0.00	***

Note: Differences between child clinics and all-generation clinics and their statistical significance are reported in the right-most column. In addition to these covariates, we control for interaction terms between municipality-level characteristics in the baseline year (i.e., 1999) and time trends. Descriptive statistics of these variables are not reported in this table because of space limitations. The association between the MSCI expansion from 1999 to 2011 and these variables are summarized in Online Appendix B.

Table 3: Monthly Number of Visits

	(1)	(2)	(3)
Child Clinic	169.641***	169.681***	-26.736
	[21.439]	[21.441]	[29.681]
Elig	-1.073	-1.105	-0.53
	[1.914]	[1.934]	[1.832]
Elig * Child Clinic	6.219**	6.207**	6.648***
	[2.440]	[2.437]	[2.314]
Public Clinic	-390.792***	-390.848***	122.081*
	[84.421]	[84.436]	[65.861]
Not-for-profit Clinic	-372.441***	-372.700***	241.097***
	[135.443]	[135.392]	[67.011]
Corporate Clinic	544.861***	544.843***	224.242***
	[13.641]	[13.644]	[21.460]
Beds	125.644***	125.595***	41.985*
	[42.854]	[42.851]	[21.645]
Population		0.000	0.000
		[0.000]	[0.000]
Share of Elderly		-156.883	-805.636
		[1,413.363]	[1,161.648]
Share of Children		648.233	703.289
		[1,302.780]	[1,063.764]
Year Fixed Effects	yes	yes	yes
Municipality Fixed Effects	yes	yes	no
Clinic Fixed Effects	no	no	yes
Prefecture-specific Trend	no	yes	yes
Municipality-level Characteristics as of 1999	no	yes	yes
Obs.	62,221	62,221	62,221
R2	0.27	0.27	0.87

Note: The sample consists of all pediatric clinics. The dependent variable is the monthly number of visits in each clinic. Standard errors are clustered at municipality level. "Municipality-Level Characteristics as of 1999" represents the interaction terms of linear time trends and municipality-level characteristics as of 1999. Details on the municipality level characteristics included in this analysis are provided in Appendix B. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

Table 4: Monthly Number of Visits: Detailed Results

	Municipality FEs			Clinic FEs		
	First	Follow-Up	Off-hour	First	Follow-Up	Off-hour
	(1)	(2)	(3)	(4)	(5)	(6)
Child Clinic	45.302***	124.379***	14.603***	-31.298**	4.562	5.506
	[6.974]	[20.969]	[3.490]	[12.421]	[29.232]	[7.788]
Elig	-0.951**	-0.153	0.475*	-1.417***	0.886	0.309
	[0.436]	[1.832]	[0.260]	[0.498]	[1.800]	[0.337]
Elig * Child Clinic	5.280***	0.927	-0.414	7.327***	-0.678	-0.103
	[0.833]	[2.512]	[0.332]	[1.071]	[2.239]	[0.515]
Year Fixed Effects	yes	yes	yes	yes	yes	yes
Municipality Fixed Effects	yes	yes	yes	no	no	no
Clinic Fixed Effects	no	no	no	yes	yes	yes
Prefecture-specific Trend	yes	yes	yes	yes	yes	yes
Municipality-level Characteristics as of 1999	yes	yes	yes	yes	yes	yes
Obs.	62,221	62,221	45,789	62,221	62,221	45,789
R2	0.15	0.26	0.04	0.67	0.86	0.46

Note: All estimates control for the same covariates as column (2) in Table 2. Dependent variables are monthly number of first visits, follow-up visits and off-hour visits. "Municipality-Level Characteristics as of 1999" represents the interaction terms of linear time trends and municipality-level characteristics as of 1999. Details on the municipality level characteristics included in this analysis are provided in Appendix B. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

Table 5: Effects on Practice Location Choice

	Full Sample			New Clinic	
	(1)	(2)	(3)	(4)	(5)
Child Clinic	132.696	118.894	1.796	-341.434*	-451.123**
	[84.331]	[85.904]	[4.256]	[177.713]	[190.865]
Elig	5.300	2.972	0.087	-23.178	-12.794
	[3.757]	[3.455]	[0.251]	[24.393]	[23.876]
Elig * Child Clinic	-13.232	-13.723	0.043	63.734***	70.435***
	[9.880]	[9.996]	[0.325]	[20.471]	[21.699]
Public Clinic		-760.417***	10.666		-610.472**
		[189.067]	[9.348]		[305.342]
Not-for-profit Clinic		308.093	8.238		48.593
		[358.310]	[8.397]		[367.904]
Corporate Clinic		-98.908**	-3.517		-15.523
		[38.561]	[2.928]		[83.902]
Beds		8.682	2.746		17.830
		[40.985]	[3.330]		[221.524]
Population		-0.001	0.000		-0.019***
		[0.001]	[0.000]		[0.005]
Share of Elderly		2,185.716	3.810		-17,611.054
		[2,853.635]	[172.379]		[17,896.834]
Share of Children		-2,476.073	176.218		6,843.008
		[1,788.721]	[233.287]		[17,478.368]
Year FEs	yes	yes	yes	yes	yes
Municipality FEs	yes	yes	no	yes	yes
Clinic FEs	no	no	yes	no	no
Prefecture Specific Trend	yes	yes	yes	yes	yes
Obs.	62,221	62,221	62,221	3,761	3,761
R2	0.889	0.889	0.990	0.886	0.888

Note: Dependent variable is population within a radius of 3 km from where the clinic is located. Population data are from the 2005 census. In columns (4) and (5), the sample consists of newly opened clinics that are defined as those first appearing in the Survey on Medical Institution but not appearing in the previous wave. "Municipality-Level Characteristics as of 1999" represents the interaction terms of linear time trends and municipality-level characteristics as of 1999. Details on the municipality level characteristics included in this analysis are provided in Appendix B. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

Table 6: Detailed Results on Practice Location

	(1)	(2)	(3)
Definition of Location Characteristics			
Population Density within 1 Km	119.793***	120.612**	119.761**
	[46.088]	[46.953]	[46.496]
Population Density within 3 Km	63.734***	65.118***	70.435***
(Baseline)	[20.471]	[21.307]	[21.699]
Population Density within 5 Km	34.687**	35.326**	41.018***
	[14.162]	[14.738]	[15.059]
Binary Outcome : Rural (1) vs Other(0)	0.001	0.001	0.001
	[0.002]	[0.002]	[0.002]
Population Density : SAA	109.540**	115.202**	112.424**
	[51.609]	[52.228]	[54.692]
Effects by Age Groups			
Population Density : SAA age $<$ 15	8.705	9.402	8.351
	[7.155]	[7.200]	[7.681]
Population Density : SAA 15 $<=$ age $<$ 65	92.129**	96.397**	95.723**
	[40.334]	[40.849]	[42.169]
Population Density : SAA age $>$ 65	20.718**	21.526**	22.486**
	[9.554]	[9.799]	[10.020]
Hospital Level Covariates	no	yes	yes
Year Fixed Effects	yes	yes	yes
Municipality Fixed Effects	yes	yes	yes
Clinic Fixed Effects	no	no	no
Prefecture-specific Trend	yes	yes	yes
Municipality-level Characteristics as of 1999	no	no	yes

Note: This table reports coefficients and standard errors of the interaction term ($Child * Elig$). Definitions of the dependent variables are found in the left-most column. "population density" is that in the SAA where each clinic is located. Population within a radius of 1, 3, and 5 km is calculated from fourth-order mesh data of population from the 2005 census. "population density" is calculated from the census data of 2000, 2005, and 2010. The sample consists of newly established clinics which are defined as those first appearing in the Survey on Medical Institution but not appearing in the previous wave. "Municipality-Level Characteristics as of 1999" represents the interaction terms of linear time trends and municipality-level characteristics as of 1999. Details on the municipality-level characteristics included in this analysis are provided in Appendix B. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

Table 7: Effects on the Number of Consultation Days

	Consultations Days (1)	Days>4 (2)	Days>5 (3)	Days>6 (4)
Panel A. Municipality FEs				
Elig * Child Clinic	-0.008***	-0.001**	-0.002	0.000
	[0.002]	[0.001]	[0.002]	[0.000]
Obs.	62,221	62,221	62,221	62,221
R2	0.14	0.11	0.19	0.04
Panel B. Clinic FEs				
Elig * Child Clinic	0.000	0.000	0.001	0.001*
	[0.004]	[0.001]	[0.001]	[0.000]
Obs.	62,221	62,221	62,221	62,221
R2	0.64	0.65	0.83	0.75
Panel C. Opening Clinics				
Elig * Child Clinic	-0.012	0.00	-0.004	-0.002
	[0.010]	[0.002]	[0.005]	[0.002]
Obs.	3,761	3,761	3,761	3,761
R2	0.21	0.22	0.24	0.15
Year Fixed Effects	yes	yes	yes	yes
Municipality Fixed Effects	yes	yes	yes	yes
Prefecture-specific Trends	yes	yes	yes	yes
Municipality-level Characteristics as of 1999	yes	yes	yes	yes

Note: All estimates control for the same covariates as for those in column (2) of Table 2. Panel A reports the results from the estimation that controls for municipality fixed effects, but Panels B controls for clinic-level fixed effects. Panel C includes only newly opened clinics and controls for municipality fixed effects. "Municipality-Level Characteristics as of 1999" represents the interaction terms of linear time trends and municipality-level characteristics as of 1999. Details on the municipality-level characteristics included in this analysis are provided in Appendix B. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

A Details on the MSCI Expansion

Table A1: Average MSCI Eligible Age by Prefecture: 1999–2011

Prefecture	1999	2002	2005	2008	2011	Response Rate
Hokkaido	1.7	3.7	6.1	6.9	7.2	77.6
Aomori	3.4	3.4	3.7	4.7	6.9	70.7
Iwate	3.5	4.9	6.0	6.0	8.2	90.6
Miyagi	1.4	1.6	3.2	3.5	5.0	94.9
Akita	3.0	6.0	6.0	6.0	6.5	88.4
Yamagata	3.5	6.0	6.0	6.0	7.6	63.5
Fukushima	2.9	6.1	6.1	7.7	11.1	83.9
Ibaraki	2.2	2.3	5.1	6.9	10.3	82.7
Tochigi	2.4	6.0	7.7	9.3	14.5	77.7
Gunma	4.8	6.1	6.4	11.3	15.0	66.2
Saitama	2.4	4.9	5.7	6.8	12.8	86.9
Chiba	2.2	2.6	3.1	6.4	10.5	91.8
Tokyo	5.2	6.0	6.7	14.6	14.6	78.2
Kanagawa	2.6	3.9	5.3	6.5	7.0	88.0
Niigata	0.7	3.1	4.9	7.7	11.0	92.0
Toyama	3.7	6.0	6.1	6.8	8.7	93.4
Ishikawa	1.9	5.4	7.0	8.7	11.3	71.5
Fukui	2.2	5.7	5.2	6.2	13.6	90.5
Yamanashi	2.6	5.3	8.0	8.2	12.6	75.3
Nagano	4.2	5.8	5.9	7.2	10.7	69.2
Gifu	2.2	3.0	4.8	10.9	14.5	80.6
Shizuoka	1.8	4.2	6.0	7.0	10.3	82.8
Aichi	2.2	3.2	5.6	9.6	13.3	96.9
Mie	2.1	2.4	3.6	5.1	7.3	68.1
Shiga	1.3	2.0	3.9	6.0	6.7	63.4
Kyoto	2.2	2.3	6.0	6.2	6.5	85.9
Osaka	1.8	2.7	4.0	5.7	7.6	69.4
Hyogo	2.1	5.2	6.0	9.1	10.4	71.9
Nara	2.0	2.0	3.1	6.0	6.6	85.3
Wakayama	2.2	2.3	4.1	6.1	6.9	68.4
Tottori	1.6	3.3	4.2	6.4	15.0	76.1
Shimane	3.3	6.0	3.6	6.1	8.3	49.8
Okayama	2.4	2.6	4.2	7.0	9.5	89.7
Hiroshima	1.5	2.5	6.0	6.5	12.8	78.8
Yamaguchi	2.6	2.6	6.0	6.1	8.6	91.0
Tokushima	2.6	2.6	4.1	7.1	10.4	35.1
Kagawa	3.3	6.0	5.6	5.8	6.2	75.8
Ehime	2.2	2.2	3.0	9.9	10.1	88.6
Kochi	1.0	2.0	3.7	6.2	6.4	68.7
Fukuoka	2.3	2.3	3.2	5.6	6.3	78.1
Saga	2.8	2.8	3.5	5.3	10.2	61.8
Nagasaki	2.4	2.4	5.7	6.0	7.4	82.7
Kumamoto	3.1	4.6	5.9	6.7	7.9	86.0
Oita	2.5	2.5	4.8	9.3	9.9	95.0
Miyazaki	2.1	4.5	4.7	4.9	10.2	85.7
Kagoshima	5.9	5.9	6.0	6.0	7.4	79.2
Okinawa	2.9	3.7	3.7	4.8	5.0	71.6

Note: Average MSCI eligible age is calculated as a population-weighted average of eligible age in municipalities within a prefecture. The eligible age of municipalities is based on an original survey conducted by the author. The weighted response rate of this survey is reported in the right column. If the eligibility criteria is set as "preschool children," a value of 6 is assigned because children start elementary school in April at the age of 6 years.

Table A2: Distribution of MSCI Eligible Age

	1999		2002		2005		2008		2011	
Eligible Age	N	Share	N	Share	N	Share	N	Share	N	Share
No MSCI	13	2%	6	1%	2	0%	0	0%	0	0%
0	15	2%	0	0%	0	0%	0	0%	0	0%
1	44	7%	9	1%		0%	1	0%	0	0%
2	242	40%	152	25%	62	10%	10	2%	2	0%
3	160	26%	138	23%	78	13%	25	4%	10	2%
4	24	4%	44	7%	39	6%	9	1%	3	0%
5	13	2%	33	5%	24	4%	9	1%	1	0%
6	96	16%	220	36%	372	61%	362	59%	208	34%
7	0	0%	0	0%	2	0%	7	1%	0	0%
8	0	0%	0	0%	2	0%	4	1%	5	1%
9	0	0%	2	0%	14	2%	49	8%	67	11%
12	1	0%	3	0%	8	1%	53	9%	89	15%
15	4	1%	5	1%	9	1%	83	14%	215	35%
18	0	0%	0	0%	0	0%	0	0%	12	2%

Note: The data are from 614 municipalities between 1999 and 2011.

B Determinants of the MSCI expansion from 1999 to 2011

To establish the conditional independence of the MSCI expansion, the choice of the municipality level covariates is particularly important. In addition to the basic covariates such as demographics, we also control for several municipality-level characteristics such as the number of child clinics as of 1999. Here, we directly check the determinants of MSCI expansion from 1999 to 2011 by estimating the following equation,

$$Elig_{m,2011} - Elig_{m,1999} = \beta_0 + \beta_1 Elig_{m,1999} + \beta_2 X_{m,1999} + Pref_p + \nu_{mt}, \tag{25}$$

where $Elig_{m,2011} - Elig_{m,1999}$ is the difference in MSCI eligible age from 1999 to 2011; $X_{m,1999}$ is a vector of municipality-level characteristics as of 1999; $Pref_p$ is prefecture fixed effects; ν_{mt} is an error term. As potential determinants of MSCI expansion, we control for demographic characteristics such as population, share of children under 15 years old, share of elderly population aged over 65 years old and share of women. Also controlled for are household characteristics (e.g., share of nuclear household and one-person household and the average number of household members). In addition, the economic environment as of 1999 is controlled for with the average taxable income per person, the industrial structure (i.e., share of secondary and tertiary workers), and the average land price. Finally, we control for the number of child clinics per 100,000 people to adjust the basic environment of pediatric care. Note that, in this analysis, municipalities that provided complete responses for the MSCI eligible age from 1999 to 2011 are included.

Table B1 summarizes the results from ordinary least squares regression. In column (1), we explore the relationship between income level and future MSCI expansion, conditional on basic demographics. The coefficient of the average taxable income is 0.869 and statistically significant, suggesting that relatively affluent municipalities expanded MSCI eligible age from 1999 to 2011. In column (2), we additionally control for prefecture fixed effects. Results in this column suggest that, with prefecture, there is no positive association between income level and MSCI expansion. In column (3), we include all the base-line characteristics of municipalities and additionally include prefecture fixed effects in column (4). In both columns, the share of nuclear household and one-person household is negatively associated with MSCI expansion. This suggests that the municipalities with a relatively larger share of small households were not likely to expand MSCI eligibility. In addition, the number of child clinics was positively associated with MSCI expansion This was probably because the municipality level association of pediatricians could often persuade their mayor to extend MSCI eligible age. Finally, average land price is positively associated with MSCI expansion, indicating that financial situation might be an important determinant of future MSCI expansion. These results again suggest that extension of MSCI eligible age *itself* is endogenous. This endogeneity may potentially violate the common trend assumption under naive difference-in-differences. Therefore, our triple differences analysis compares the child clinics and all-generation clinics and controlled potential effects from municipality-level characteristics by the interaction terms (time trends and municipality characteristics) as of 1999.

Table B1: Determinants of MSCI Expansion From 1999 to 2011

Characteristics of the Municipalities in 1999	(1)	(2)	(3)	(4)
Eligible Age	-0.657***	-0.853***	-0.698***	-0.903***
	[0.075]	[0.080]	[0.081]	[0.085]
Population	-0.000***	-0.000***	-0.000***	-0.000***
	[0.000]	[0.000]	[0.000]	[0.000]
Child Share	-25.786**	17.315	-34.665***	-4.63
	[11.247]	[11.092]	[11.873]	[13.946]
Elderly Share	3.279	24.263***	-0.165	3.92
	[6.838]	[5.890]	[8.757]	[7.616]
Number of Pediatric Clinics			1.217	-0.723
			[7.822]	[6.575]
Share of Females			-16.700***	-5.332
			[3.654]	[3.797]
Population Density			0	0
			[0.000]	[0.000]
Share of Nuclear Households			-11.388***	-10.141***
			[2.501]	[2.779]
Average Household Size			-0.082	0.044
			[0.163]	[0.139]
Share of One-person Households			-17.724***	-11.192***
			[3.212]	[3.369]
Average Taxable Income	0.869**	-0.059	0.742*	0.317
	[0.361]	[0.313]	[0.409]	[0.414]
Share of Secondary Workers			-9.305	-21.258**
			[7.591]	[8.521]
Share of Tertiary Workers			-12.624	-21.101**
			[7.790]	[8.567]
Land Price			0.000***	0
			[0.000]	[0.000]
Const.	10.935***	5.497*	51.447***	44.668***
	[3.417]	[2.864]	[8.843]	[10.721]
Obs.	614	614	614	614
R2	0.121	0.538	0.318	0.573
Prefecture Fixed Effects	no	yes	no	yes

Note: Fitting by ordinary least squares. The dependent variable is the difference in MSCI eligible age from 1999 to 2011. In columns (2) and (4), prefecture fixed effects are controlled for. In this analysis, municipalities that provided complete responses on the MSCI eligible age from 1999 to 2011 are included. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

C On the Potential Endogeneity of $Child_{it}$

One concern identified in this paper is that the choice of official specialty *itself* may also be affected by regional unobservables. Although specialty choice of primary care physicians is strongly affected by education received in medical university, physicians freely choose their official specialties in Japan. For example, pediatricians who mainly learned specialized care for children at their medical school generally want to provide primary care only for children, but if the competitive pressure is very strong, they may unwillingly decide to profess "internal medicine" in addition to "pediatrics". For these pediatricians, the MSCI expansion may have favorable effects because, under a generous MSCI, they can earn sufficient income even without additionally professing "internal medicine". These possibilities may cast a concern that $Child_{it}$ is no longer exogenous for the MSCI expansion.

In addition to potential endogeneity of $Child_{it}$ from flexible specialty choice, $Child_{it}$ can be affected by $Elig_{it}$ if child or all-generation clinics move to municipalities with a generous MSCI. While inter-municipality migration of pediatricians is not likely because municipality-level medical associations hold strong power over the new entry of primary care physicians, it is also possible that a generous MSCI attracts more pediatricians.

To check these possibilities, we examine whether the MSCI eligible age has statistically significant effects on specialty choice. At first, we estimate the impacts of MSCI eligible age on the choice of "pediatrics" among all clinics. Then, we explore the MSCI effects on the choice of being a child clinic among all pediatric clinics. If child clinics move to municipalities with a generous MSCI, we may find positive effects from the MSCI. Table C1 summarizes the results. In columns (1) and (2), we explore the association between MSCI eligibility and the choice of "pediatrics" and being a child clinic, respectively. Note that some unobservable municipality level trends violate the common trend assumption here because we exploit naive difference-in-differences in this analysis, but we find there are no statistically significant effects on specialty choice from MSCI eligible age. This may suggest that the endogeneity of $Child_{it}$ in our main specification is not so severe.

Next, we check the endogeneity of $Child_{it}$ *via* another specification. As explained previously, $Child_{it}$ can be endogenous to the MSCI expansion if child or all-generation clinics move to municipalities with a generous MSCI. To check this point directly, we count the number of child and all-generation clinics in each municipality and survey year, and examine how the numbers are associated with MSCI eligible age. By using the balanced panel data of 611 municipalities for 5 years (n = 3,055), we estimate following equation,

$$N_{mt} = \theta_0 + \theta_1 Elig_{mt} + \theta_2 Z_{mt} + \theta_3 T * Z_{m,1999} + \theta_4 Pref_p + \varepsilon, \tag{26}$$

where N_{mt} is the number of child/all-generation clinics in a municipality m in year t. As in the specification in the main analysis, we control for municipality level demographics (Z_{mt}) as well as the interaction term of time trend and municipality-level characteristics as of 1999 ($T * Z_{m,1999}$). Results are

Table C1: MSCI Expansion and Specialty Choice

	(1) Ped	(2) Child
Elig	-0.001	0.000
	[0.000]	[0.001]
Government-owned Clinics	0.077***	0.039
	[0.026]	[0.035]
Other Public Clinics	-0.230***	-0.042
	[0.026]	[0.107]
Cooperative Clinics	0.003	0.026***
	[0.003]	[0.007]
Beds	-0.023***	-0.030***
	[0.006]	[0.010]
Population	0.000	-0.000***
	[0.000]	[0.000]
Share of Children	-0.279	-1.010*
	[0.432]	[0.591]
Share of Elderly	1.048**	0.422
	[0.423]	[0.499]
Sample	All Clinics	Pediatric Clinics
Year Fixed Effects	yes	yes
Municipality Fixed Effects	yes	yes
Prefecture-specific Trends	yes	yes
Linear Trends of City Characteristics	yes	yes
Obs.	235028	62221
R2	0.128	0.198
Mean of Dependent Variable	0.265	0.184

Note: Fitting by ordinary least squares. In columns (1) and (3), The dependent variable is a binary variable that takes a value of 1 for "pediatric clinics" whose official specialties include "pediatrics". In columns (1) and (3), the dependent variable is a binary variable that takes a value of 1 for "child clinics" whose official specialties include "pediatrics" but not "internal medicine". Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

summarized in Table C2. The dependent variable is the number of child clinics in columns (1) and (2) and the number of all-generation clinics in columns (3) and (4), respectively. In columns (2) and (4), we report the subsample results for the South Kanto region. All results from columns (1) to (4) show that the number of child/all-generation clinics is not associated with the MSCI eligible age, suggesting that inter-municipality migration of clinics is not likely to our empirical analysis.

Table C2: MSCI Expansion and the Number of Clinics

	Child Clinic		All-Gen Clinic	
	All	Kanto	All	Kanto
	(1)	(2)	(3)	(4)
Elig	0.01	0.01	0.01	0.06
	(0.012)	(0.054)	(0.042)	(0.164)
Population	0.000***	0.000***	0.00	0.00
	0.000	0.000	0.000	0.000
Share of Child	6.598	72.890***	-15.73	167.60
	(4.776)	(18.310)	(10.854)	(73.221)
Share of Elderly	10.637**	45.106***	56.971**	68.02
	(4.445)	(15.717)	(15.997)	(40.526)
Year Fixed Effects	yes	yes	yes	yes
Municipality Fixed Effects	yes	yes	yes	yes
Prefecture-sSpecific Trends	yes	yes	yes	yes
Municipality-level Characteristics as of 1999	yes	yes	yes	yes
Observations	3,055	550	3,055	550
N. of Municipality	611	110	611	110
R-squared	0.60	0.72	0.83	0.85
Summary of Dep.				
Mean of Dep.	4	7	21	38
SD	11	15	50	59
min	0	0	0	0
max	154	154	835	548

Note: Fitting by ordinary least squares. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

D Detailed Results on Consultation Days

In addition to the results on the number of consultation days, days on which physicians do not have consultations is also investigated. In Table D2, we create binary variables that takes 1 if a clinic is open on Monday (column 1) and Sunday (column 7). Except for the results on Thursday in column (4) and Sunday in column (7), we find negative and statistically significant effects on the MSCI expansion. For example, the increase of the MSCI eligible age by 15 years decreases the likelihood of clinics conducting consultations on Wednesday by 4.5 percentage points and in Saturday by 3 percentage points, respectively. [34] Finally, we explore the effects on opening hours by am, pm, and overtime in Table D1. The dependent variable in columns (1) and (3) is the number of days when a clinic open in the am, pm, and overtime, which ranges from 0 days to 7 days. Here, we find a statistically significant reduction in the number of days when a clinic is open during overtime hours. Given that physicians choose to reduce their labor supply from the hours when they feel dissatisfaction, this is an intuitive result.

Table D1: Opening Hours

	AM (1)	PM (2)	Overtime (3)
Child Clinic	0.129***	0.332***	-0.129**
	[0.026]	[0.042]	[0.065]
Elig	0.004	-0.014	0.011
	[0.004]	[0.012]	[0.009]
Elig * Child Clinic	-0.005**	-0.002	-0.020***
	[0.003]	[0.004]	[0.007]
Year Fixed Effects	yes	yes	yes
Municipality Fixed Effects	yes	yes	yes
Prefecture-sSpecific Trends	yes	yes	yes
Municipality-level Characteristics as of 1999	yes	yes	yes
Obs.	62,221	62,221	62,221
R2	0.14	0.24	0.28

Note: All covariates that are included in Table D2 are controlled for. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

[34] Note that we find statistically significant positive effects on Sunday. This is a counterintuitive result because we find opposite results for other days. However, the irregular result on Sunday is probably due to the reduced out-of-pocket expenditure being generally higher on weekends under the same MSCI system because the national government set additional copayment for visits on weekends. If patients are more likely to visit clinics on weekends as a result of MSCI expansion, some pediatricians may choose to open their clinics on Sunday.

Table D2: Effects on the Number of Consultation Days According to Day of the Week

	Mon	Tue	We	Thu	Fri	Sat	Sun
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Child Clinic	0.021***	0.021***	0.020**	0.014	0.022***	0.030***	-0.006
	[0.004]	[0.005]	[0.009]	[0.010]	[0.004]	[0.006]	[0.006]
Elig	0.001	0.001	0.001	-0.001	0.001	0.001	0.000
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.000]
Elig * Child Clinic	-0.001**	-0.001*	-0.003**	-0.001	-0.001***	-0.002***	0.001*
	[0.000]	[0.001]	[0.001]	[0.001]	[0.000]	[0.001]	[0.001]
Public Clinic	-0.206***	-0.223***	-0.160***	-0.129***	-0.202***	-0.702***	0.074***
	[0.027]	[0.027]	[0.028]	[0.032]	[0.026]	[0.031]	[0.021]
Not-for-profit Clinic	-0.306**	-0.152**	-0.277**	-0.271**	-0.181**	-0.810***	-0.043**
	[0.126]	[0.077]	[0.131]	[0.132]	[0.090]	[0.110]	[0.018]
Corporate Clinic	0.002	0.005*	0.012**	0.026***	0.001	0.003	0.019***
	[0.002]	[0.003]	[0.005]	[0.006]	[0.002]	[0.003]	[0.004]
Beds	0.010***	0.013***	0.022***	0.030***	0.010***	0.020***	0.008
	[0.003]	[0.003]	[0.006]	[0.007]	[0.003]	[0.004]	[0.006]
Population	-0.000**	-0.000**	-0.000*	0.000	-0.000**	-0.000***	0.000
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Share of Elderly	-0.410	-0.765	0.481	-1.225*	-0.704	-1.457***	-0.539
	[0.491]	[0.473]	[0.559]	[0.652]	[0.452]	[0.520]	[0.441]
Share of Children	-0.389	-0.612*	0.196	-0.285	-0.492	-1.167***	-0.399
	[0.342]	[0.360]	[0.443]	[0.618]	[0.344]	[0.439]	[0.356]
Year Fixed Effects	yes	yes	yes	yes	yes	yes	yes
Municipality Fixed Effects	yes	yes	yes	yes	yes	yes	yes
Prefecture-specific Trends	yes	yes	yes	yes	yes	yes	yes
Municipality-level Characteristics as of 1999	yes	yes	yes	yes	yes	yes	yes
Means of Dependent Variables	0.96	0.95	0.89	0.82	0.96	0.92	0.05
Observations	62,221	62,221	62,221	62,221	62,221	62,221	62,221
R Squared	0.10	0.08	0.08	0.12	0.08	0.20	0.06

Note: Dependent variables take 1 if a clinic accepts patients in a given day of a week. For clinic ownership, the reference variable is clinics owned by individual physicians. "Municipality-level Characteristics as of 1999" represents the interaction terms of linear time trends and municipality-level characteristics as of 1999. Details on the municipality-level characteristics included in this analysis are provided in Appendix B. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

E Heterogeneous Effects

E.1 Clinic Characteristics

In summary, the MSCI expansion increases the number of patients treated in clinics and accelerates the concentration of pediatricians into densely populated areas. In addition, physicians who are affected by the MSCI expansion generally reduce their consultation days. These results are sufficiently robust to the changes in empirical specification, as is shown in the previous section.

To supplement these main findings, this section provides subsample analysis according to the ownership of clinics. In general, clinics in Japan can be divided into two types. One type of clinic is owned by solo-practice physicians. Although the management of these individual clinics can be supported by neither other clinics through business alignment nor governmental transfer, they have large discretion in choosing their practice location and consultation days. Therefore, we expect that their choice of practice location and consultation days is responsive to the MSCI expansion. On the other hand, clinics that are owned by medical corporations and public clinics may not respond to the MSCI expansion because they do not have such discretion. Although these two types of clinics face similar increases in patient demand, their responses may be very different.

In Panel A of Table E1, we show the heterogeneous effects by creating the interaction term of a binary variable for clinics owned by solo-practice physicians ($solo_{it}$) and the triple-differences term ($Child_{it} * Elig_{mt}$). Here, we find that the point estimate in the interaction term with solo-practice physicians is positive but not statistically significant. For the number of follow-up visits, we find large heterogeneity in the impact; point estimate of the interaction term is 6.713 and statistically significant. Because clinics owned by solo-practice physicians treat lower numbers of patients in general, these results may suggest that impact on the number of visits is much larger for these clinics. For location choice and consultation days, we find no particular heterogeneities in columns (4) and (5).

In Panel B of Table E1, we create an interaction term of clinics with beds ($Beds_{it}$) and the triple-differences term. Although the heterogeneous effects are not precise for most covariates, point estimates are generally negative and large. For example, on the number of visits, the coefficient of the interaction term is -10.719, which is larger than that of the triple-differences term (8.095) in the absolute term. This suggests that clinics with beds, which generally treat children with chronic conditions such as asthma, are not so affected by the MSCI expansion. In addition, the coefficient of the interaction term also takes a large negative value (-868.721) in column (4), suggesting that effects on the location on this subsample is almost null.

E.2 Regions

In Table E2, we split the sample by South Kanto region and Other because the South Kanto region, which includes Tokyo, is the most populated area in Japan.[35] Results on the South Kanto region are summarized in Panel A. In this region, we find strong effects on the total number of visits in columns (1) to (3) and the

[35]The South Kanto region consists of Tokyo, Saitama, Chiba, and Kanagawa prefectures.

number of consultation days in column (5). However, the effects on the practice location choice is negative and not statistically significant in column (4), suggesting that there are no positive effects on this outcome. On the other hand, we find the opposite results in other regions in Panel B. In Panel B, the effects on the number of visits and consultation days are generally weak and not statistically significant, but we find that positive statistically significant effects on the population within a radius of 3 km in column (4). The null effects on the practice location choice in the South Kanto region is probably because of the high land price. It is well known that land prices in the South Kanto region, especially Toyko, are high, and it is therefore possible that additional revenue from MSCI is not enough to be located in densely populated areas.

Table E1: Heterogeneous Effects by Ownership

	Visits			Pop 3km	Consultation
	Total	First	Re		days
	(1)	(2)	(3)	(4)	(5)
Panel A. Clinics Owned by Solo-practice Physician					
Elig * Child Clinics	5.136	6.969***	-1.833	1,870.682**	-0.006**
	[3.395]	[1.034]	[3.528]	[756.227]	[0.003]
Elig * Child Clinics * Solo	4.203	-2.509***	6.713**	119.661	0.000
	[3.484]	[0.909]	[3.388]	[465.010]	[0.003]
Panel B. Clinics with Beds					
Elig * Child Clinics	8.095***	5.699***	2.396	1,987.594***	-0.006**
	[2.461]	[0.839]	[2.502]	[624.761]	[0.003]
Elig * Child Clinics * Beds	-10.709	-3.392*	-7.317	-868.721	-0.001
	[8.397]	[2.006]	[7.691]	[1,182.972]	[0.005]
Year Fixed Effects	yes	yes	yes	yes	yes
Municipality Fixed Effects	yes	yes	yes	yes	yes
Prefecture-specific Trends	yes	yes	yes	yes	yes
Municipality-level Characteristics as of 1999	yes	yes	yes	yes	yes

Note: Heterogeneous effects on the clinics owned by solo-practice physicians are reported in Panel A. Those on the clinics with beds are reported in Panel B. In all estimations, clinical specialty, ownership, and demographic variables are additionally controlled for. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

Table E2: Heterogeneous Effects by Regions

	Visits			Pop 3 km	Consultation
	Total	First	Re		days
	(1)	(2)	(3)	(4)	(5)
Panel A. South Kanto					
Elig * Child Clinics	10.839***	4.322***	6.517*	-534.265	-0.009**
	[3.332]	[0.881]	[3.491]	[392.510]	[0.004]
Obs	19,915	19,915	19,915	19,915	19,915
Panel B. Other Regions					
Elig * Child Clinics	4.485	7.876***	-3.39	340.043*	-0.004
	[3.313]	[1.135]	[3.106]	[201.940]	[0.004]
Obs	40,869	40,869	40,869	40,869	40,869
Year Fixed Errors	yes	yes	yes	yes	yes
Municipality Fixed Errors	yes	yes	yes	yes	yes
Prefecture-specific Trends	yes	yes	yes	yes	yes
Municipality-level Characteristics as of 1999	yes	yes	yes	yes	yes

Note: South Kanto regions include Tokyo, Kanagawa, Saitama, and Chiba prefectures. In all estimation, clinical specialty, ownership, and demographic variables are additionally controlled for. Standard errors are clustered at the municipality level. ***, $p < 0.01$. **, $p < 0.05$, *, $p < 0.1$.

F Proofs in the Theoretical Framework

F.1 Proofs of Lemma 1

Because $w > B$ and $\delta_k \geq \delta_l$ if $N_k > N_l$, from (14) and (16), $P_k^* < P_l^*$ if $N_k > N_l$. Therefore, the following condition must be satisfied to satisfy (21).

$$p_h \frac{H_k}{D_k} - c < p_h \frac{H_l}{D_l} - c, \tag{27}$$

which leads to $H_k/D_k < H_l/D_l$. In the full-coverage case, from (17),

$$\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)B]\frac{N_k}{D_k} < \frac{\beta}{\theta p_h}[\delta_k w + (1-\delta_k)B]\frac{N_k}{D_k} = \frac{H_k}{D_k} < \frac{H_l}{D_l} = \frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)B]\frac{N_l}{D_l}, \tag{28}$$

which leads to $D_k/N_k > D_l/N_l$ if $N_k > N_l$. In the partial-coverage case, from (18),

$$\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)\theta B]\frac{N_k}{D_k} < \frac{\beta}{\theta p_h}[\delta_k w + (1-\delta_k)\theta B]\frac{N_k}{D_k} = \frac{H_k}{D_k} < \frac{H_l}{D_l} = \frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)\theta B]\frac{N_l}{D_l}, \tag{29}$$

which leads to $D_k/N_k > D_l/N_l$ if $N_k > N_l$.

F.2 Proof of Proposition 1

Assume that the population is larger in area k than in area l ($N_k > N_l$). From (14) and (16), P_k^* and P_j^* are constant regardless of the coverage of public health insurance. Therefore, from 21, we obtain

$$\frac{p_h \frac{\bar{H}_k}{\bar{D}_k} - c}{p_h \frac{\bar{H}_k}{\bar{D}_l} - c} = \frac{p_h \frac{\hat{H}_k}{\hat{D}_k} - c}{p_h \frac{\hat{H}_k}{\hat{D}_l} - c}, \tag{30}$$

where $\bar{D}_k$ and $\bar{D}_l$ ($\hat{D}_k$ and $\hat{D}_l$) denote the number of physicians in the equilibrium in the full-coverage case (the partial-coverage case). From (17), (18), and (30), we obtain

$$\frac{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)B]\frac{N_k}{\bar{D}_k} - c}{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)B]\frac{N_l}{\bar{D}_l} - c} = \frac{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)\theta B]\frac{N_k}{\hat{D}_k} - c}{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)\theta B]\frac{N_l}{\hat{D}_l} - c}. \tag{31}$$

Moreover,

$$\frac{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)B]\frac{N_k}{\bar{D}_k} - c}{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)B]\frac{N_l}{\bar{D}_l} - c} - \frac{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)\theta B]\frac{N_k}{\hat{D}_k} - c}{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)\theta B]\frac{N_l}{\hat{D}_l} - c} = \tag{32}$$

$$\frac{\frac{1-\theta}{\theta}\beta B\left[(1-\delta_l)\frac{N_l}{\bar{D}_l} - (1-\delta_k)\frac{N_k}{\bar{D}_k}\right]}{\left\{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)B]\frac{N_l}{\bar{D}_l} - c\right\}\left\{\frac{\beta}{\theta p_h}[\delta_l w + (1-\delta_l)\theta B]\frac{N_l}{\bar{D}_l} - c\right\}} c > 0.$$

Therefore, $\bar{D}_k < \hat{D}_k$ and $\bar{D}_l > \hat{D}_l$ to satisfy (31).

F.3 Proof of Proposition 2

Assume that the population is larger in area k than in area l ($N_k > N_l$). From (14) and (16), P_k^* and P_j^* are constant regardless of the co-payment rate. Therefore, from (21),

$$\frac{p_h \frac{H_k}{D_k} - c}{p_h \frac{H_k}{D_l} - c} \tag{33}$$

is unchanged in the equilibrium.

$$\frac{d}{d\theta}\left(\frac{p_h \frac{\bar{H}_k}{\bar{D}_k} - c}{p_h \frac{\bar{H}_k}{\bar{D}_l} - c}\right) = \frac{\frac{p_h}{\theta}\left(\frac{\bar{H}_k}{\bar{D}_k} - \frac{\bar{H}_l}{\bar{D}_l}\right)c}{\left(p_h \frac{\bar{H}_k}{\bar{D}_l} - c\right)^2} < 0, \tag{34}$$

because $H_k/D_k < H_l/D_l$ (which is derived in the proof of Lemma 1). Therefore, $\bar{D}_k$ ($\bar{D}_k$) increases with a decrease in θ to keep (33) at a constant value. Moreover,

$$\frac{d}{d\theta}\left(\frac{p_h \frac{\hat{H}_k}{\hat{D}_k} - c}{p_h \frac{\hat{H}_k}{\hat{D}_l} - c}\right) = \frac{\delta_l \frac{\hat{N}_l}{\hat{D}_l}\left(\beta B \frac{\hat{N}_k}{\hat{D}_k} - c\right) - \delta_k \frac{\hat{N}_k}{\hat{D}_k}\left(\beta B \frac{\hat{N}_l}{\hat{D}_l} - c\right)}{\left(p_h \frac{\hat{H}_k}{\hat{D}_l} - c\right)^2} \tag{35}$$

$$< \frac{\delta_l \left(\frac{\hat{N}_l}{\hat{D}_l} - \frac{\hat{N}_k}{\hat{D}_k}\right)c}{\left(p_h \frac{\hat{H}_k}{\hat{D}_l} - c\right)^2} < 0, \tag{36}$$

because $N_k/D_k < N_l/D_l$ from Lemma 1. Therefore, $\hat{D}_k$ ($\hat{D}_k$) increases with a decrease in θ to keep (33) at a constant value.