Medical Charges and Provider Ownership in Urban China: An Empirical Analysis

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ABSTRACT

We use data from a 2005 World Bank household survey to analyze the price differential between state and private providers in urban China. We also investigate the impact of an expansion of the private sector on the price of health care. Our analysis finds strong evidence that the state providers charge higher prices for their services than the private ones. In addition, the price premium is much greater for the disadvantaged groups than for the general population, indicating that the structure of the public-private price gap is highly regressive. On the issue of market liberalization, we find consistent cost reductions associated with a large private share of the market for medical personnel. Since the overall charges and the price differential both highly elastic to the private market share, we conclude that increasing the competition in the physician market may help significantly lower the cost of health care in urban China.

KEYWORDS: Price comparison; Provider ownership; Market liberalization; Disadvantaged social groups; Semiparametric switching regression model

JEL CODES: C24, D40, I11

1. Introduction

Defining appropriate roles of the government and the market in the health sector is a difficult issue for nations throughout the world. The challenge is particularly acute for transitional countries, where the introduction of market forces has become an important development tool for the other parts of the economy. As one of the fastest growing transitional economies, China represents an intriguing case. While market liberalization in the last three decades has spurred unprecedented economic growth at large, how to reform its predominantly state-owned health sector remains to this day a gaping issue.

And this is an issue that demands urgent attention, since China's health system is undergoing a crisis. As the country devotes an increasing amount of resources to health care[1], its level of population health has stagnated[2] and popular dissatisfaction with the performance of the system mounted. A common Chinese saying goes: "*Kan bin nan, kan bin gui.*" It translates into English as: "It is difficult ("*nan*") and expensive ("*gui*") to seek medical care."

In this paper, we focus on the issue of "*gui.*"[3] Since the mid-1990s, the growth rate of medical expenditure has consistently outstripped those of real income and consumer price index.[4] In the wake of this cost escalation, however, the share of organized financing, be it from the government or an insurer, has decreased, resulting in large burdens of out-of-pocket payment for households.[5]

As in other countries that have experienced similar health cost inflation, a debate is raging on in China about the causes of this spending crisis. Although scholars and policy-makers generally agree on certain issues, such as insufficient public financing and distorted provider incentives[6], there is much less consensus on the effects of provider

[1]The number of doctors per thousand people increased from 1.17 in 1980 to 1.54 in 2007. During the same period, health spending as a share of GDP rose from 3.15% to 4.67%. [Ministry of Health (2008)]

[2]Several key health status measures have experienced a "regression to the mean." While they were well above the level expected of a country with low per capita income before the reform era, health measures such as life expectancy, infant mortality and under-five mortality rates, have regressed to the world average by 2000, despite rapid economic growth. [Wang (2003), Wagstaff (2004), Eggleston et al. (2005)]

[3]Manifestations of the problem of "*nan*" vary by geography. While scarcity of health resources forces many living in rural areas to travel to the city for even simple medical conditions, long queuing and waiting time at a hospital is a daily nuisance for most urban residents.

[4]Inflation-adjusted per-episode outpatient cost grew by 13.29% per annum between 1996 and 2003. The same growth pattern was observed for inpatient expenditure. In contrast, the annual growth rate of real income during the same period was 8.9% and 2.5% in urban and rural areas, respectively. [Center for Health Statistics and Information, Ministry of Health (2004)]

[5]At the turn of the millennium, private financing accounted for 60% of China's total health spending, a dramatic increase from the level of 21.2% in 1980. By 2006, the figure had slightly slid to 52.2%, thanks mostly to the gradual expansion of various insurance schemes. [Ministry of Health (2008)] A study by van Doorslaer et al. (2005) found that households in China were much more likely to experience "catastrophic" health expenses (defined as expenditure that exceeds 40% of household non-food consumption) than elsewhere in Asia.

[6]For instance, see Liu and Hsiao (1995), Eggleston and Yip (2004), Meng and Tang (2004), Liu et

ownership and market competition.

One group of researchers argues that the state sector behaves like a conglomerate of collusive oligopolists. Instead of pursuing public health goals, state-owned providers abuse their predominant market status, especially in urban areas, and rely on government policies to obstruct private entry. Therefore, they have been able to charge exorbitant prices with impunity. The natural solution is to break down state dominance by encouraging competition from private providers. Another line of argument insists that the reason for public providers' profit-seeking behavior is that two decades of health care reform have introduced an overly strong economic incentive to the health sector. Increasing market competition with private providers, most of whom are profit oriented, will only sanction and hence intensify the motive to over-supply and overcharge medical services. The solution prescribed by this school is to remove the impact of economic interests by nationalizing the health sector, at least at the level of basic care. Competition from private providers is allowed only for certain high-end services.[7]

At the center of the debate are two issues essential to the efficiency of any health system: provider ownership and competition. Because of its peculiarities, health care does not lend itself to obvious conclusions on either front. Unlike for other sectors, economic theory does not give definitive predictions about how provider ownership or market liberalization influences health care outcomes and the overall performance of the health system. The vast empirical literature that has burgeoned over the years offers mixed evidence on these issues. Furthermore, this literature focuses heavily on developed countries, such as the US. Empirical evidence from developing and transitional economies is in severe short supply. [Sloan (2000)] Differences in socioeconomic priorities and institutional characteristics make it difficult for results obtained from one type of countries to find direct applications in the other.

In this paper, we aim to alleviate this shortage in the literature by investigating two issues in the context of China: (1) How do state and private health sectors compare with one another in an important aspect: the cost of care, as measured by the price charged for the services delivered? (2) Is there an association between the cost of care and the market share of private providers? The answers we provider here will inform the health care reforms in other developing and transitional economies, with whom the Chinese health system shares many salient features[8]. Moreover, our analysis will shed light on how characteristics unique to the health system under study shape answers to the two universal questions regarding provider ownership and market competition.

To avoid future confusion, we pause here for a moment to make an important distinction between the *price of individual services* and the *price of health care*. In China, the former is regulated by government in the form of official fee schedules. By contrast, the

al. (2000), Liu et al. (2003), World Bank (2005).

[7]Another rationale for the "pro-government" argument is the public goods property of health care.

[8]In terms of the supply side of the system, common features include dominance of state ownership and limited administrative capacity. [Eggleston et al. (2005)]

latter is jointly determined by patient and physician decisions. It depends on not only the fees, but also the type and quantity of care delivered. Therefore, to treat *the same condition of the same level of severity*, the price of care may still differ substantially from case to case depending on factors such as preferences and the style of practice. In this paper, we are interested in the price of health care, since it represents the real cost to the household in its consumption of medical services.

The data we use come from a household survey the World Bank conducted in five metropolitan areas in December 2005. The survey provides information on actual expenditure on the most recent outpatient visit, ownership of the provider used and a set of patient/provider characteristics. To study the first question shown above, we define the public-private price differential as the difference in average expenditure between the state and the private health sectors. A key contribution of our paper is that it uses an endogenous switching regression model to correct for sample selection bias in the patient-level data. The model is estimated with both parametric and semiparametric techniques. Specific attention is paid to the comparison between the results of these two approaches. Semiparametric estimation is robust to the violation of distributional assumptions on the error terms. Compared to the parametric model, therefore, it affords stronger confidence in our assessment of the performance of the two health sectors in urban China.

To answer the second question, we use a series of simple linear regressions to show the association between a set of market share measures for private providers and (1) the actual expenditure; (2) the estimated public-private price differential. A negative sign in front the market share regressor indicates that a strong private sector presence not only decreases the average cost of care, but also reduces the price gap between the public and the private sectors. Since the coefficient of interest belongs to an aggregate-level variable estimated using micro data, we use cluster sample methods to obtain correct inference.

The main results of the paper are as follows. While the numeric value of the price gap is rather sensitive to the smoothing parameters chosen in the preferred semiparametric estimation, the general picture is that outpatient visits are significantly more expensive at public hospitals than at private clinics, even after controlling for unobserved factors related to patient choice. Moreover, the gap is larger for vulnerable social groups, including the poor, the uninsured and rural migrants, than for the general population. It may be the case that the poor are constrained by a low ability to pay to choosing low skill providers, so there is a fundamental difference in the quality of care between the two health sectors. However, to the extent that differences in provider characteristics are observed by patients and taken into account in their selection, our model sufficiently controls for these unobserved heterogeneities. Therefore, a more plausible explanation is differing abilities to price discriminate. In light of the domination of the urban health care market by state-owned providers[9], this presents a substantial barrier to access for

[9]As will be shown in Section 2, state-owned institutions possess a predominant share of the market thanks largely to legal barriers to entry and insufficient information on provider performance.

the socially disadvantaged. In addition, we find in the structural selection equation that patients' choice of provider by ownership is insensitive to the public-private price differential. Given the relatively shallow insurance coverage in China, in terms of both its width and depth[10], this result suggests the importance of non-price factors, such as hospital location and perception of service quality, in China's health care demand function. Finally, we find consistent negative correlations between the average cost of care and private shares in the market for physicians, especially those with advanced medical degrees. The same is true for the public-private price gap. Reduction in the price differential indicates a convergence of pricing levels between the two sectors. Since the average cost is also negatively correlated with private market shares, it is plausible that the more expensive state providers converge toward the cheaper private sector as the share of doctors working for non-government hospitals increases.

How we interpret the above results has important policy implications. The first question is whether the public-private price gap should be of concern to policy-makers. Obtained in a model of individual choices, the price gap seems to be the result of utility-maximizing decisions. Nonetheless, there is still something government can do to increase social welfare. As typical for health care, the choice of provider is usually based on incomplete information. In particular, lack of quality information about private providers may prevent effective selection on price. Therefore, the price gap can be thought of as a "premium to market status." In addition, since the gap is larger for disadvantaged social groups, government may want to address the issue out of equity concerns. Policies that aim to facilitate the growth of low-cost private institutions or encourage discriminating pricing on the basis of the patient's socioeconomic status may considerably increase social welfare. To the extent that the apparently low cost consciousness is due to weak consumer bargaining power and an inability to judge the quality of services, the international literature suggests that the advantage of the insurer in cost containment, through such measures as selective contracting and provider payment method, should be emphasized in designing the health insurance system. Furthermore, the government or, in the case of China, regional insurance funds can play an important role in the collection, organization and dissemination of quality-related information in order to alleviate the informational asymmetry that may have obstructed effective competition on price among health care providers in urban China. 2nd part result suggests an effective market-based approach is to liberalize personnel market.

The remainder of the paper is organized as follows. Section 2 describes the institutional characteristics of China's urban health care system. Section 3 reviews the literature on provider ownership and competition in the health care market. In section 4, we describe the data set. Sections 5 and 6 examine the two main issues of interest: the public-private price gap, and the impact of market liberalization on the price of care. Finally, section 7 concludes.

[10]In our sample, 32% of the patients are uninsured. Among the insured, the average reimbursement rate is 37%.

2. Institutional Background

China's urban health delivery system is organized as a three tier system, consisting from bottom to top of street level primary care facilities, secondary level hospitals and tertiary hospitals. The level of technical capability rises along the ladder. The system is characterized by state ownership. Up to the mid-1990s, almost all hospitals were public non-profit organizations. Since the mid-1990s, several municipalities have experimented with introducing private hospitals. These institutions are usually much smaller than their non-profit counterparts. In 2003, for example, non-profit hospitals employed only 4% of the beds and 3% of the physicians working in urban hospitals. [China Health Yearbook (2004)] However, the private sector has grown more rapidly than the public sector. Between 1996 and 2001, the number of outpatient visits paid to state-owned facilities grew by only 2.1% while it increased by 300% at the private sector. [Chinese Health Statistical Digest (2003)]

Concurrent with the growth of private providers, the Chinese government has reduced its financial support for the state-owned health sector as a share of the total operational cost. Government subsidies now cover only basic wages and the cost of equipment maintenance. In fact, state-owned hospitals have a high degree of financial autonomy as they enjoy complete discretion over the use of income surplus. The bulk of their income comes from user charges, which are paid on a fee-for-service (FFS) basis according to a regulated fee schedule. Technically, the FFS method reimburses the hospital for each medical service provided. Since it rewards the provider for the quantity of care, this method is usually associated with high levels of use and even over-use of profitable services.

Such is indeed the case in China. The fees of many items on its regulated schedule were originally set below cost to ensure wide access to care. However, as public hospitals found their user charges falling short of recovering the operational costs, the authorities allowed them to have handsome mark-ups on pharmaceutical products and high-tech diagnostic tests as a means of cross subsidization. [Liu et al. (2000)] The result is widespread practice of excessive prescriptions of drugs and diagnostic tests. For instance, a study of village clinics found that prescriptions were strongly associated with the need to generate revenue, rather than with the patients' clinical conditions. [Zheng et al. (2003)] Consequently, China has one of the world's highest shares of drug costs in total medical costs, standing at 54.7% and 44.7% in 2003 of outpatient and inpatient expenditure, respectively. [China Health Yearbook (2004)] The average of the OECD countries is about 15%.

The above description applies mainly to the state health sector, while the policies governing the non-state sector differ somewhat. In the case of price regulation, for example, private non-profit institutions are also subject to a mandated fee schedule, but with some flexibility within a range around the official fee. Charges by private for-profit facilities, on the other hand, are not regulated. Neither type of private provider receives government subsidy, while the non-profit institutions are exempt from income taxes.

[Eggelston et al. (2007)]

On the demand side of the health care equation, insurance coverage has declined in urban areas. The share of insured population dropped from over 70% in the early 1990s to below 60% in 2003.[11] [China Health Yearbook (2005)] The drop in the spread of insurance is attributable to changes in the economic structure. Before the economic reform started, the urban health insurance system had two pillars, the government insurance scheme (GIS) and the labor insurance scheme (LIS). The former covered government employees, soldiers, students and teachers, while the latter workers at state-owned enterprises. As the reform goes on, an increasing section of the economy becomes independent of the state while the private sector grows rapidly. As a result, many people have fallen out of the old publicly based medical safety net. In an effort to expand insurance coverage in the urban areas[12], in the late 1990s, the government consolidated the GIS and LIS into a single insurance program for employees, called the urban workers' basic medical insurance (BMI) system. The program is currently undergoing another expansion, to the entire urban population[13], including the unemployed, the self-employed and those with unstable employment.[14]

When a patient, be it insured or uninsured, makes a contact with the medical system, (s)he chooses the hospital and doctor to seek care from. In other words, there is no gate-keeping system in urban China. Without credible quality information, patients usually prefer the ones that have the best reputation and market standing. These are often large state-owned tertiary hospitals, contributing to the crowdedness and over-use at these facilities. Although price differentials do exist between different levels of hospitals in both the regulated fee schedule and amount of insurance reimbursement, they are not large enough to remedy the problem.

3. Literature Review

• Ownership

Evidence on the impact of provider ownership in the market for health care comes mostly from the U.S., where the differences between for-profit and non-profit hospitals have been extensively analyzed. Some studies suggest that ownership and profit status do not make a difference in the performance of the hospitals. [e.g., Sloan (2000)] On the

[11]The vast majority of rural residents, some 80%, are uninsured, although the rate declined somewhat between 1998 and 2003 because of a new insurance initiative.

[12]Another important motivation was to increase the level of risk pooling from individual "work units," as was the case with both GIS and LIS, to the city level.

[13]This does not include migrants from rural areas. Most of them do not have official residential status in cities and are hence not entitled to local benefits.

[14]In July 2007, the State Council passed the decision to launch the urban resident BMI program. According to the official schedule, pilot experiments were carried out in 79 localities in 2007, coverage is to expand to 50% of the target population in 2008, to 80% in 2009 and finally to universal coverage by 2010.

contrary, others have found that for-profit providers have lower technical quality and higher mortality rates. [e.g., Devereaux et al. (2002)]

The limited evidence available from China is rather mixed. Some analysts have argued that the public and the private for-profit providers do not differ significantly in their behavior. They are just as likely to induce unnecessary demand or, if properly paid, to deliver preventive care. [Meng et al. (2000)] In addition, the mortality rates at private providers do not differ statistically from those of similar government hospitals. [Eggelston et al. (2007)]

At the same time, other analysts have indicated that the state sector charges a higher price for its services. In a study that is directly related to ours, a research team in Wenzhou (of Zhejiang Province) has compared the charges of public and private hospitals for a number of diseases that involved hospitalization. They have found that the state sector invariably had higher charges and longer lengths of stay. [Research Team, Wenzhou Health Economics Association (2004)] However, the study has failed to adjust for the case-mix of each hospital, thus subjecting their results to biases that may arise from patient heterogeneity. Our analysis will borrow from the literature of sample selection to address this issue.

• Competition

The literature on competition in the health care market is again based mostly on the U.S. experience. It has documented that, prior to the advent of managed care, hospitals competed in quality to attract patients covered by generous indemnity insurance. This led to a rapid spread of advanced technologies, a phenomenon referred to as "*medical arms race*" [Salkever 1978; Robinson and Luft (1985), Dranove et al. (1993)]. During this period, competition not only led to higher medical prices, but also was socially wasteful. [Dranove and Statterthwaite (2000)]

In the mid-1980s, the rise of managed care and the change in the provider payment method[15] transformed the locus of competition from quality to price. Some studies of this period show that hospital competition may not only lower the cost but also improve the outcome of medical care. [Kessler and McClellan (2000)]

A natural question in the context of our study is: which type of competition is more likely to take place in China's hospital care market, where many people pay for health care out of pocket? In addition, for reasons to be elaborated later, our analysis focuses only on outpatient care, for which price is probably a more important factor than for hospitalization. Given these conditions, competition in price seems a more plausible outcome.

However, the answer to this question is not straightforward and health economists have come up with various explanations for why competition may not have the de-

[15] In particular, the change from fee-for-service to prospective payment.

sired effect. These include, among others, imprecision of information on price and quality [Dranove and Satterthewaite 1992], search cost [Satterthewaite 1979], physician-induced demand [Fuchs 1978] and brand loyalty [Grabowski and Vernon 1992]. Therefore, whether more private entry leads to lower prices is an empirical question that needs to be closely examined.

3. Data

Data come from a household survey that the World Bank administered in December 2005 in preparation for its country report on China. The survey was conducted in five large cities: Shanghai (Pudong area), Shenzhen, Dalian, Xi'an and Chengdu. In each city, a multi-stage sampling method was used to draw random samples of households. First, communities were selected from the central urban areas with a probability-proportional-to-size (PPS) method. Then within each community, a systematic sampling strategy was employed to draw households. In total, the sample includes 4,802 households (or 15,009 individuals), of which about 30% are immigrant families.

The dataset contains cross-sectional observations on the utilization of health care services. The data fields essential to our study include: individual and family characteristics, provider characteristics and expenditure on the most recent episode of care. We recoded some of the variables on the basis of the raw data. The survey question on provider ownership, the key variable of our interest, included five categories: state-owned, private domestic, private joint-venture, private foreign and others. We merged the latter four categories into one and defined it as the "private sector." Definitions of all the other variables used in the analysis are provided in Appendix 1.

In the second part of the analysis, we supplement the World Bank data with regional socio-economic indicators collected from statistics yearbooks.

4. Public-Private Price Differential

□ **Methodology.** To estimate the price differential, we compare the average charges at the state and the private sectors, conditional on a set of individual and provider characteristics. Given the substantial financial burden that inpatient care in urban China places upon patients, especially those from the poor households, the degree of simultaneity between income, one of the regressors, and expenditure, the dependent variable, may be high in the inpatient setting. Recognizing this possibility, we use only outpatient data in our study.

We split the sample of individuals that reported on the most recent outpatient visit into two sub-samples, one using the state and the other the private sector. The public-private price differential is thus the difference in the average costs of the two sectors. Since patients choose which type of provider to seek care from, the assignment to the

two sub-samples is not random. In particular, we allow for the patient's choice to be determined by two factors: (1) the expected cost difference and (2) unobserved (to the researcher) factors that influence selection, such as preferences and the severity of illness.

This is a general specification that lies between Roy's (1951) model, where selection is based purely on the expected outcome, and Heckman's (1979, 1990) sample selection model,[16] where unobserved factors alone lead to non-random assignment. We adopt this specification because, apart from the size of the public-private price difference, we are also interested in examining the strength of its influence on the patients' choice of providers.

We start the model by specifying the cost of care in each sector. Let y_{si} and y_{pi} denote log-scale expenditure on an outpatient visit at a state and a private provider, respectively. Note that for each individual, only the cost in one sector is observed. In addition, let $\mathbf{x}_i\beta_s$ and $\mathbf{x}_i\beta_p$ be the conditional expectations of the expenditures, where $\mathbf{x}_i$ is a vector of observable patient and provider characteristics. It also includes four city dummies to control for regional heterogeneities in the price of care. β_s and β_p are unknown parameters. We can then write the reduced-form cost functions as:

$$y_{si} = \beta_{0s} + \mathbf{x}_i\beta_s + \varepsilon_{si}, \tag{1}$$

and

$$y_{pi} = \beta_{0p} + \mathbf{x}_i\beta_p + \varepsilon_{pi}, \tag{2}$$

where ε_{si} and ε_{pi} both have zero mean and are independent of $\mathbf{x}_i$.

The model is completed by defining the selection equation. Let s_i^* be a latent variable that measures the net utility of individual i receiving treatment at a state hospital. It is a function of the public-private cost difference, the "value added" offered by the state sector, or $\mathbf{z}_i\gamma$, and random shocks. The variables $\mathbf{z}_i$ are exogenous characteristics that may overlap with $\mathbf{x}_i$, and γ is a vector of parameters to be estimated. Then, assuming that individuals maximize the net utility, we can write the decision equation as:

$$s_i^* = \alpha(y_{si} - y_{pi}) + \mathbf{z}_i\gamma + \varepsilon_{0i}. \tag{3}$$

[16] Heckman's original model was specified for the case of incidentally truncated data, where only the outcome of non-censored sub-sample was observed. For an application of Heckman's model to the estimation of average treatment effects, see Green (1997), pp. 981-982.

Following the standard estimation methods, the variance of ε_{0i} is normalized to 1. When $s_i^* > 0$, the individual chooses a state provider. Otherwise, we observe a choice of a private facility. Since higher cost reduces utility, the parameter associated with the expected cost difference, α, is expected to be negative.

Equations (1)-(3) form the so-called endogenous switching regression model. [Jones (2000)] Identification of the model is achieved through imposing exclusion restrictions. In particular, we identify the structural selection equation with a dummy variable that indicates whether unnecessary care was used. This variable appears in $\mathbf{x}$, but not $\mathbf{z}$. The use of unnecessary care significantly increases expenditure. However, sufficient anecdotes show that in popular perception, public hospitals are as likely as private ones to over-prescribe. Therefore, the anticipation of unnecessary care does not influence the choice of state or private providers. To identify the cost functions, we use the ownership of the patient's employer, which is a binary variable that takes the value of 1 if the employer is owned by government. The rationale is that, under the old government and labor insurance schemes,[17] health expenditure at an appointed state hospital was fully reimbursed. Because of this legacy, employees of the state sector have a stronger tendency than the general population to use public providers, even in the new insurance system. One potential problem of using employer ownership as the identifying variable is that it could directly influence the spending on care, if the state sector offered more generous benefits than the other employers. However, such is not the case currently in China, where the new BMI system provides equal benefits across all sectors.

In the traditional parametric approach to the estimation of the model, the error terms $(\varepsilon_0, \varepsilon_s, \varepsilon_p)$ are assumed to be trivariate normally distributed. The effect of unobserved heterogeneity on the outcome equations assumes a specific functional form; i.e. the inverse Mill's ratio, which is a known function of the probability of selecting the treatment.[18] However, regression diagnostics following Pagan and Vella (1989) reject the joint normality assumption. Moreover, we find evidence of substantial heteroskedasticity in both outcome equations. Therefore, we adopt a generalized estimation framework that allows the terms that capture the expectations of ε_s and ε_p conditional on hospital ownership to enter the cost equations nonparametrically. In this manner, not only the covariance structure but also the correlation between the choice and the cost equations are left completely unspecified.

In particular, estimation proceeds in three steps. First, we substitute equations (1) and (2) into equation (3) to derive a reduced-form selection function. Since the normality assumption was not rejected for the choice function, this is estimated as a probit model. Second, constraining the coefficients associated with the city dummies to be equal across the provider sectors,[19] we combine equations (1) and (2) into a fully inter-

[17]These were the main insurance programs in urban China before being replaced by the Basic Medical Insurance (BMI) system in 1998. The former covered civil servants while the latter insured employees of state-owned enterprises.

[18]For detailed descriptions of the parametric approach, see Lee (1979), Vella and Verbeek (1999), Jones (2000) and Wooldridge (2001).

[19]This is equivalent to restricting the inter-city differences in expenditure to be equal across the

acted model, written as follows:

$$y_i = \beta_0+\beta_1 s_i+\mathbf{x}_i\beta_2+(\mathbf{x}_i*s_i)\beta_3+\rho_1(s_i*K_s(Z_i,S_i)+\rho_2[(1-s_i)*K_p(Z_i,S_i)]+\mathbf{CITY}_i\phi+\eta_i \quad (4)$$

where s_i equals 1, if the provider is state-owned, and 0, otherwise. $K_s(Z_i,S_i) = E(\varepsilon_{si} \mid Z_i, S_i = 1)$ and $K_p(Z_i,S_i) = E(\varepsilon_{pi} \mid Z_i, S_i = 0)$.[20] The vector $\mathbf{CITY}_i$ contains four city dummies. Finally, the error term η_i has zero mean and is uncorrelated with the regressors.

Equation (4) is a partial linear model, where the functional form of K_s and K_p is unknown. A well-known result in the treatment effects literature provides useful simplification for estimation. Heckman (1980) and Heckman, Ichimura, Smith and Todd (1998) demonstrate that, under the assumption of index sufficiency, which is the standard assumption for laten index models, the error terms in the outcome equations depend on the Z_i only through the probability of receiving treatment. In other words, $K_{s/p}(\cdot)$ in equation (4) are functions of only variable, $\hat{P}$, the probability of choosing a public hospital. With this useful simplification, we employ the simple differencing-based estimation method developed by Yatchew (1997, 1998, 2003) for partial linear models. Yatchew showed that this approach is valid only if the number of arguments in the nonparametric function is no greater than 3.

The differencing-based estimation proceeds in various steps. First, the data is sorted, *within each sub-sample*, by ascending values of the probability of selecting a public hospital estimated from the first-stage probit regression. Second, an m-order differencing with optimal weights is taken of the reordered data. For example, if the order is 1, the outcome equations become

$$Y_{mi} - Y_{m,i-1} = [X_i - X_{i-1}]\beta_m + [K_m(\hat{P}_i) - K_m(\hat{P}_{i-1})] + [\varepsilon_{mi} - \varepsilon_{m,i-1}],$$

where $m = s$ *or* p.

Yatchew shows that, if K_m are continuous and have bounded derivatives, the 2nd square bracket of the expression disappears in large samples. Then simple OLS can be run on the transformed data to obtain consistent and $\sqrt{n}$-normal estimates of β_m. In the context of this paper, we run OLS regression on equation (4) using the transformed data, without the $K(\cdot)$ functions. We experimented with differencing orders of 1, 5, 10, 25 and 50 in estimation.[21] Yatchew (2003) gives the optimal differencing weights and

public and the private sectors.

[20] These would be the inverse Mill's ratios in the parametric model.

[21] In general, as the order of differencing increases, the estimator approaches asymptotic efficiency.

the method to adjust for heteroskedasticity.

Before estimating equation (4), it is necessary to define a common support for $\hat{P}$ between the state and the private samples. Heckman, Ichimura, Smith and Todd (1998) showed that estimating the average treatment effects outside of the common support introduces a bias arising from comparing the incomparable. In the trimmed data, about 79% of the public ($S = 1$) and 89% of the private samples ($S = 0$) lie in the overlapping region.

Notice that the differencing method absorbs the intercept terms in the cost equations. Since intercepts are essential in the computation of the average treatment effect, we use the method developed by Andrews and Schafgans (1998) to recover these terms. In particular, the estimator is

$$\hat{\beta}_{0m} = \frac{\sum_{i=1}^{Nm}(y_i - \mathbf{x}_i\hat{\beta}_m)S_i d(\mathbf{z}_i\gamma - \lambda_m)}{\sum_{i=1}^{Nm} S_i d(\mathbf{z}_i\gamma - \lambda_m)},$$

where $m = s \; or \; p$. λ_m is a parameter whose value is determined by the researcher. Called the bandwidth or smoothing parameter, it is chosen so that $\lambda_s \to \infty$ as $N_s \to \infty$ and $\lambda_p \to 0$ as $N_p \to \infty$. Finally, the function $d(\cdot)$ for the state sample is given by

$$\begin{aligned} d(x) &= 1 - \exp(-\frac{x}{b-x}), \quad for\ x \in (0, b); \\ &= 0, \quad for\ x \leq 0; \\ &= 1, \quad for\ x \geq b. \end{aligned}$$

Following Andrews and Schafgans (1998), b is set equal to 1. The function for the private sample is very similar except the negative sign within the exponential function disappears. The function $d(\cdot)$ can be thought of as a weighting scheme that assigns larger weights to the observations greater (smaller) than and far away from the threshold λ in the public (private) sample than to the others. Essentially, λ is used to to control the proportion of observations used in the estimation of the intercept. The intuition is that the intercept is identified at those points where the probability of selecting a public hospital is close to 1 (or 0 for the private sample) and is 1 (or 0) in the limit as $N \to \infty$. This is an example of "identification at infinity." In the estimation, we experimented with using 2.5%, 5%, 10% and 20% of the observations.

Finally, $X\hat{\beta}_m$ are moved to the left side of equations (1) and (2) to create a new dependent variable, $\widetilde{y}_m = y_m - X\hat{\beta}_m$. Then various nonparametric smoothing methods can be used to estimate the equation

$$\widetilde{y}_m = K_m(\hat{P}) + \varepsilon_m.$$

According to Yatchew, this modular approach yields consistent and asymptotically normal estimates of the nonparametric functions, since the convergence rate of the estimated coefficients $\hat{\beta}_m$ is much faster than that for $K_m(\hat{P})$.

After the estimation of the partial linear model, we compute the predicted values of y_{si} and y_{pi} from (4), and insert them back into equation (3) to obtain consistent (but inefficient) estimates of the parameters of the structural selection equation. Since we do not adjust for the fact that the regressors in this stage contain estimates, the standard errors will be underestimated.[22]

This step also produces the estimated price differential, which is defined as the sample average of the predicted cost difference. In particular, it is equal to:

$$CD = \hat{\beta}_1 + \bar{\mathbf{x}}\hat{\beta}_3 \tag{5}$$

where $\bar{\mathbf{x}}$ is the sample mean of the regressors, and $\hat{\beta}_1$ and $\hat{\beta}_3$ are estimated parameters.

Finally, we single out sub-groups of patients defined by their income, employment, insurance and migrant status. Then by varying the sub-sample upon which the average cost difference is calculated, we also estimate the price differential relevant for the different social groups.

The cost difference given in (5) is on the log scale. However, what we are really interested in is the price differential in unlogged, monetary units. Therefore, we use the (heteroskedasticity-robust) retransformation method described by Ai and Norton (2006) to convert the log difference into the raw difference. Their paper also shows how to obtain standard errors of the retransformed value.

□ **Results.** Since occupational characteristics are necessary for identification, individuals under the age of 19 are excluded. We also screened out cases treated at specialized facilities, such as hospitals for communicable diseases. The reason is that specialized care represents an area where a high degree of differentiation exists between the public and the private sectors. Facilities for disease prevention are usually government-owned, since the low profitability of these services is unlikely to attract private providers. Moreover, it has been documented that specialization in services not reimbursable by the urban health insurance program (BMI) is a strategy frequently used

[22]This turns out not to be a problem since, as will be shown later, the unadjusted estimate of α, the parameter of real interest in the structural selection equation, is statistically insignificant. Further adjustment will only reduce its significance level.

by private providers to circumvent the dominant public sector. Given the differentiation, the specialized hospitals must be excluded to increase the level of comparability between the two sectors. Finally, after dropping the cases with missing values, we obtain a sample of 1,572 observations, of which 84% were treated at the state sector.

Table 1 shows the descriptive statistics of the key variables for the state and the private sub-samples. As the table demonstrates, the average charge for an outpatient visit is higher at public facilities. The patient population treated at the state sector is on average older, wealthier, better insured and more likely to be local residents. Moreover, a public provider is more frequently chosen for specialist consultations.

In specifying the model, it has been assumed that the patient's family income, employment, migrant status and possession of non-job based insurance are exogenous to both the choice of a provider and the expenditure on health care. In reality, our observations of these variables are the results of individual choices that may well depend on unobserved factors. Fortunately, the survey does report self-evaluated health status, which allows us to control for the most important factor that drives these choices.[23]

Results of the first-stage probit model are omitted.[24] The ownership of the patient's employer, the identifying variable for the cost functions, is found to be a strong predictor of the choice of a public hospital. The point estimate is 0.37, which is significant at the 1% level. Diagnostic tests of normality and heteroskedasticity found no misspecification of the model.

Since the results with higher orders of differencing appear to be sensitive to the choice of the smoothing parameter λ,[25] we report only the results obtained with 10 order of differencing and where 10% of the observations in each sub-sample are used to estimate the intercept terms. Table 2 presents the coefficient estimates from the linear regression of equation (4). The standard errors are robust to heteroskedasticity in the cost equations. To demonstrate the importance of correcting for self-selection, the results of an OLS regression without the selectivity terms are also presented.

Consistently, the estimates obtained with the OLS model are very similar to those under the switching regression. Selectivity seems to have very modest impact on individual and provider characteristics. The only appreciable effect it has on the result is through changing the coefficient before the provider ownership variable. According to the uncorrected OLS, as is shown in the row indicated as "State-owned" of the table, the average charge for a baseline patient is fairly similar between the two sectors. Once the selectivity is controlled for, however, an outpatient visit becomes much more expensive at public than at private hospitals. This finding indicates that the unobserved factors

[23]The data does not allow us to address state-dependent reporting biases. Despite this apparent limitation, we rest this issue in peace by noting the widespread use of self-evaluated health status in health economics research.

[24]They are available from the authors upon request.

[25]To our knowledge, there is no study that examines the robustness of the differencing-based estimator to the order of differening.

that influence the choice between the state and the private sectors lead to a "lump-sum" increase in the cost of care for all patients, regardless of their (observable) characteristics.

Some other findings in Table 2 also warrant discussion.[26] We note the rather substantial public-private difference in the individuals socio-economic characteristics. In particular, poor patients spend considerably less than the wealthy in private facilities, *ceteris paribus*. This pattern is also observed in the state sector. However, the gap in spending becomes much smaller. This finding suggests a rather unpleasant aspect of China's urban health care system: health care at the state sector brings tremendous financial burdens for the disadvantaged. Contrary to social expectations, public hospitals are far from the advocates for the vulnerable social groups.[27]

We also tested the partial linear model with two alternatives: (1) the impact of the probability of selecting a public hospital on the cost of care is insignificant; (2) the $K(\cdot)$ functions take the form of inverse Mill's ratios. Using the specification tests described by Yatchew (2003), both null hypotheses were rejected at the 5% significance level. In other words, if uncontrolled for, self-selection would have caused significant bias in the analysis. Moreover, joint normality of the error terms is an inappropriate assumption given our data.

Since we are essentially interested in the public-private difference in the average charge of care *on the raw scale*, we must convert the cost difference defined in equation (5). Since the test based on Pagan and Vella (1989) finds strong evidence of heteroskedasticity in the cost functions, we use the heteroskedasticity-robust retransformation methods described in Ai and Norton (2006). In particular, let Y_s and Y_p denote the raw-scale expenditures in the state and the private sectors, respectively. Then the price ratio can be written as:

$$\frac{Y_s}{Y_p} = \exp(CD) \times \exp(\sigma_s^2(\mathbf{x}) - \sigma_p^2(\mathbf{x})) \tag{6}$$

where $\sigma_s^2(\cdot)$ and $\sigma_p^2(\cdot)$ are variance functions.

Then the price differential is equal to the expression in equation (6) minus 1, and then multiplied by 100%. We also calculate the retransformed price differential for various sub-samples. The results are shown in Table 3.

Our estimation produces strong evidence that, with the selectivity bias corrected, the

[26]It must be recognized that the estimated coefficients in Table 1 are not the marginal effects of the variables on the conditional expectation of outpatient cost. To arrive at these marginal effects requires two additional steps. First, for variables that also appear in the selection equation, we must remove their impact on the selection from the coefficient. [Green (1997)] Second, we must calculate their marginal effect on the retransformed cost variables and the relevant standard errors. [Ai and Norton (2006)] We skip both steps since the parameters *per se* are not our primary interests.

[27]Note that the model we use has controlled for the patients' health level and utilization patterns.

average charge for an outpatient visit is higher at the state sector than at the private facilities. As Table 3 shows, the estimated price difference is 98%. If we consider that part of the price gap that is attributable to patient selection, both observed and unobserved, as legitimate, this "unexplained" cost difference suggests that health care in state-owned hospitals is indeed more expensive. Moreover, by choosing different sub-samples based on income, insurance and migrant status, with the exception of employment, we show that this price premium is much greater for the disadvantaged groups than for the general population. These results are consistent with our earlier discussion of the estimated coefficients associated with the socio-economic indicators. Based on these results, we conclude that the structure of the public-private price gap in urban China is highly regressive.

Striking as the results may appear, they are hardly improbable in China, where the common perception is that services at the state sector are excessively expensive [Lim *et al.* (2004)]. A plausible explanation for the regressive price gap is that managers/physicians at the private facilities are less burdened with non-market oriented objectives than their public colleagues. They are also less subject to the price regulation in health care. According to the current regulations, for-profit private providers have complete pricing autonomy, while non-profit private hospitals enjoy some degree of freedom within a wide interval around the official fee schedule. Public providers must follow the mandated price regime. Therefore, it is much easier for private providers to engage in price discrimination, varying the charges for the same services according to the patients' ability to pay.

Finally, the third-stage probit estimation of the structural selection equation shows that the expected cost difference does not have a significant effect on the patients' choice of provider type.[28] This finding is somewhat surprising, since, given the large share of out-of-pocket payments in health care spending in China, one would expect the price differential to have a significant impact on cost-conscious consumers. One probable explanation for our paradoxical result is that the provider-consumer informational asymmetry is so severe as to blunt the patients' ability to "shop for prices." This hypothesis also explains the dominance of the health market in urban China by the expensive tertiary-level public hospitals.

5. Expanding Private Participation in the Health Care Market

□ **Methodology.** In this part of the analysis, we examine the association between the market share of the private providers and the average cost of outpatient care. Operationally, this is equivalent to asking:"*With the individual characteristics properly controlled for*, to what extent can the regional variation in average cost be attributed to differences in the market share of private providers?"

[28]The point estimate of the coefficient, α, is -0.68, with a 95% confidence interval of (-2.66, 1.31).

By adding regional variables, including the market share indicator, to equation (4), we could estimate the impact of the private market share on y_i. As Wooldridge (2003) shows, however, the standard errors obtained with this naive method are biased, overstating the significance level of the regional variables.

To make correct inference, we follow the two-step procedure introduced in Wooldridge (2003). First, we use the city dummies to obtain estimates of the five city-specific intercepts, δ_g. Second, we substitute these estimates into the equation

$$\delta_g \quad = \quad \theta_0 + \theta_1 m_g + \mathbf{w}_g \theta_2, \qquad\qquad g = 1, ..., 5$$

where m_g is the market share indicator and $\mathbf{w}_g$ is a vector of other regional variables.

If the number of elements in $(m_g, \mathbf{w}_g)$ equals 4, we obtain a unique solution for the parameters; i.e., $\hat{\theta} = (\hat{\theta}_0, \hat{\theta}_1, \hat{\theta}_2) = W^{-1}\hat{\delta}$, where W is the 5×5 matrix with the gth row defined as $(1,\ m_g, \mathbf{w}_g)$. Wooldridge (2003) shows that, as long as the assumptions necessary for the consistency of the OLS regression are satisfied, the two-step estimator of θ is consistent. It is asymptotically normal with the variance $X^{-1}V_{\hat{\delta}}(X^{-1})'$, where $V_{\hat{\delta}}$ is the variance-covariance matrix of $\hat{\delta}$.[29]

In the same spirit, we can also examine the association between the average price gap in a city and the private market share in its health care sector. A negative value of the coefficient before the private market share indicator, θ_1, will suggest that a strong private sector is associated with both lower average costs of health care and a smaller price differential.

Since the number of cities is extremely small, only limited room is available to address inter-city heterogeneities and other specification issues. To overcome this restriction, we run repeated estimations for each sample, alternating the market share indicator and the other controls used. This can be thought of as a rough robust test for our model specification. Table 4 shows the control variables that appear in the estimations. Choosing one variable from each column and running the estimation on the two samples separately will yield a total of 320 procedures.

□ **Results.** The result of the second part of the analysis varies substantially with the market share indicator chosen. When the private market share is measured by the number of physicians or outpatient visits, about 65% of the estimations based on the full sample produce a negative market-share coefficient. The percentage rises to over 80% if the number of physicians with advanced medical degrees is used. Consistent with this pattern, the median cost elasticity with respect to the private market share

[29] If $G > K + 1$, inferences in the same spirit can be made based on the minimum-distance estimator from Wooldridge (2001).

ranges from -3 in the case of outpatient visits to -5.76 for well-educated physicians.[30] A similar observation is made also of the results based on the price gap.

These findings warrant some discussion. As we demonstrated earlier, patients are not responsive to price differences because of informational asymmetry. As a result, competition for individual patients may not create enough pressure for the price of medical services to decrease. Indeed, studies of the other health care markets have suggested that competition for insurance contracts and production inputs (e.g., physicians and hospital managers) may offer greater opportunities than competing for patients, in reducing the cost of health care. [Kessler and McClellan (2000), Docteur and Oxley (2003), World Bank (2004)]

Our results are consistent with the international literature. It demonstrates that the *mechanism* of market liberalization matters. Increasing the share of outpatient visits flowing to the private hospitals seems to have only modest impact on the price of care. On the contrary, competition for physicians, the most critical input for the production of health care, achieves the most consistent and significant cost reductions. This suggests that liberalizing the market for medical personnel has the greatest potential for being translated into real competitive power for the private providers.[31]

6. Discussion

In this paper, we analyze the public-private price gap in China's urban health care market. After correcting for selection bias, we find that outpatient services are much more expensive in the state sector. Moreover, the price gap is highly regressive in that it affects disadvantaged social groups more severely than the well-off. In the second part of the analysis, we examine the effect of expanding the private health care sector. Our results show that a stronger private sector does seem influence the behavior of the state providers. Moreover, the the mechanism of market liberalization matters. In particular, liberalizing the market for medical personnel has the greatest potential for generating real competitive power from private providers.

Admittedly, our analysis may raise more questions than answers, leaving considerable room for future research. The data available does not provide direct information on market competition. The share of such things as beds, doctors and visits owned by private hospitals does not reveal the real structure of the market. We do not know whether expanded private entry has fostered head-to-head competition with the state

[30] As the distributions of the estimation results are skewed, we use the median, rather than the mean, of the distribution as the representative value.

[31] Our analysis shows that the other inputs, namely, hospital beds and medical appliances, have ambiguous impact on the average price of care. The cost elasticity with respect to the private share of beds is positive, while that associated with medical equipment is negative but very small. This finding indicates that increasing the private participation in the market for non-critical inputs may fail to bring real competitive pressure onto the public sector.

establishment. Nor are we able to tell how the process has influenced the way in which patients choose hospitals. Finally, there is no information on the outcomes of treatment, such as rate of recovery and quality of care, except for how much the patient spent. Future research can yield promising results in areas such as the impact of competition, or more precisely, private entry, on the selection process. Moreover, price should not be the sole focus of interest. The other determinants of patients' decision-making, such as perception of quality and loyalty, should receive equal, if not more, attention.

Even so, our study does reveal an interesting pattern about the link between private entry and price. In particular, the different associations between the price and various market-share indicators offer a hopeful glimpse about the most effective way to foster a competitive market for hospital care in urban China.

A final remark concerns the sustainability of the price gap. In an ideal market, where information is complete and the consequence of bad quality is moderate, demand and supply would eliminate any possibility of price gaps. Health care, however, is quite different. After all, with medical services, the medium of consumption is the patient's own body, which, once damaged, may not be replaceable. For these reasons, a price difference may exist for some types of health care services.

Another possible reason, one that is rich with Chinese characteristics, is that most private hospitals are excluded from the network of urban basic medical insurance, now the largest third-party payer in China. Then price would be higher at public hospitals simply because their services are covered by insurance.

Whatever the reason for the price gap, it is our belief, or at least hope, that it is not sustainable in the long run and that gradually competition will equate the prices between the two sectors. To find out the optimal way to achieve that calls for vigorous further research based on more informative data sets.

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Table 1. Definitions and Distributions of Key Variables

Variable	Definition	State N = 1,673 Mean(SD)	Private N = 282 Mean(SD)
Log Spending		5.53(1.05)	4.92(1.22)
Ind. Characteristics			
Male	Dummy	0.42(0.49)	0.40(0.49)
Married	Dummy	0.86(0.34)	0.89(0.32)
Self-Reported Health	Scale of 1-5. Recoded into 5 levels in estimation.	3.26(0.89)	3.34(0.80)
Migrant	Dummy, 1 for households with more than 2 members of non-local residential status	0.29(0.45)	0.53(0.50)
Family Income	Recoded into within-city quintiles in estimation.	3,332(3,117)	2,738(3,811)
Insurance			
Job-based	Dummy, 1 for BMI and GIS	0.65(0.48)	0.37(0.48)
Voluntary	Dummy, 1 for commercial insurance and medical assistance	0.07(0.25)	0.05(0.22)
Uninsured (reference)	Dummy	0.28(0.46)	0.57(0.49)
Employment			
Unemployed	Dummy	0.10(0.31)	0.12(0.33)
Retired	Dummy	0.27(0.45)	0.18(0.38)
Not Working	Dummy, 1 for non-working age	0.03(0.17)	0.01(0.12)
Employed (reference)	Dummy	0.59(0.49)	0.68(0.46)
Education			
No Education (reference)	Dummy	0.02(0.15)	0.02(0.13)
Primary Schooling	Dummy	0.10(0.31)	0.17(0.38)
Junior High	Dummy	0.29(0.45)	0.36(0.48)
High School	Dummy	0.42(0.49)	0.36(0.48)

Continues

Appendix 1: (*Continued*)

Variable	Definition	State N = 1,673 Mean(SD)	Private N = 282 Mean(SD)
Higher Education	Dummy	0.16(0.37)	0.09(0.29)
Type of Visit			
Ordinary	Dummy, 1 for ordinary visits and check-ups	0.70(0.46)	0.82(0.38)
Specialist	Dummy, 1 for specialist consultation	0.27(0.44)	0.08(0.28)
Other (reference)	Dummy	0.02(0.16)	0.10(0.29)
Prov. Characteristics			
Hi-Level Gen. Hosp.	Dummy, 1 for general hospitals at or above city level	0.81(0.39)	0.29(0.45)
Lo-Level Gen. Hosp.	Dummy, 1 for general hospitals below city level	0.17(0.38)	0.20(0.40)
Hi-Level OP Dept.	Dummy, 1 for OP departments at or above city level	0.01(0.11)	0.01(0.08)
Lo-Level OP Dept. (reference)	Dummy, 1 for OP departments below city level	0.14(0.35)	0.68(0.47)
Identifying Variables			
Selection Function			
Age	Number of years	45.62(14.38)	41.58(12.87)
Cost Functions			
State Sector	Dummy, 1 for state-owned employer	0.47(0.50)	0.25(0.43)

Table 2. Linear Regression Results

Variable	OLS[†]		Switching Regression	
	Private	State	Private	State
Constant	4.53***	5.13***	–	–
Age	0.01	0.01***	0.01	0.005**
Male	-0.16	0.04	-0.16	0.03
Married	0.04	-0.05	-0.04	0.04
Health Status				
Poor	0.62	0.68**	0.70	0.74***
Fair	1.22***	0.69**	1.21***	0.69***
Good	0.61**	0.39	0.70**	0.42***
Very Good	0.23	0.30*	0.26	0.31***
Migrant	-0.01	0.09	0.03	0.10
Employment				
Unemployed	-0.01	0.02	-0.10	0.02
Retired	-0.15	-0.07	-0.13	-0.06
Not Working	0.81**	0.27*	0.71***	0.23
Family Income				
Poorest 1/5	-0.82***	-0.23*	-0.90***	-0.26***
Next Poorest	-0.58**	-0.16*	-0.66***	-0.19**
Mid-Income	-0.74**	-0.10*	-0.75***	-0.11
Next Richest	-0.81**	-0.08	-0.85***	-0.09
Education				
Primary School	-0.39	-0.10	-0.37	-0.08
Junior High	-0.50	-0.31	-0.53	-0.31*
High School	-0.50	-0.30	-0.51	-0.30*
Higher Education	-0.85	-0.39	-0.85	-0.37**
Insurance				
Job-based	0.18	-0.01	0.11	-0.04
Voluntary	0.04	-0.10	-0.04	-0.13
Type of Visit				
Ordinary	0.58	-0.40	0.48**	-0.46***
Specialist	1.06***	-0.06	0.94***	-0.23
Type of Provider				
Hi-Level Gen. Hosp.	0.73**	0.69***	0.28	0.50***
Lo-Level Gen. Hosp.	0.70*	0.45***	0.57***	0.35***
Hi-Level OP Dept.	-0.26	0.24	-0.53	0.07
Shanghai	-0.05		-0.15	
Dalian	-0.01		-0.01	

Continues

Appendix 2: (*Continued*)

Variable	OLS†	Switching Regression
Xi'an	-0.26***	-0.21**
Chengdu	-0.49***	-0.48***
State-owned	0.91	1.68***

†: Inference based on standard errors adjusted for heteroskedasticity and within-city clustering.

Table 3: Public-Private Price Gap

Ave.	Income[a]		Residence		Employment		Insurance	
	Poor	*Rich*	*Mig.*	*Local*	*UnEmp.*	*Emp.*	*Unins.*	*Ins.*
98% (2.65)	–	–	–	–	–	–	–	–
–	**158%** (2.89)	2.1% (1.57)	**146%** (3.38)	75% (2.37)	72% (2.90)	142% (2.79)	**148%** (3.35)	67% (2.33)

[a]: The comparison groups are individuals whose household income is in the lowest and the highest quintile, respectively, of each city's income distribution.
Standard errors are in the parentheses.

Table 4: Control Variables

Private Market Share	Income	Health Resources	Population Health
Outpatient Visits	Disposable Income	Beds/1,000†	Infant Mortality
Medical Equipment	Savings Rate	Physicians/1,000†	Maternal Mortality
Beds	–	General Occupancy Rate	ER Death Rate
Physicians	–	State Occupancy Rate	ICU Death Rate
High-Edu Physicians*	–	–	–

*: These are physicians with doctoral medical degrees.
†: The number of beds/physicians per 1,000 persons.