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Would Allowing Privately Funded Health Care Reduce Public Waiting Time? Theory and Empirical Evidence from Canadian Joint Replacement Surgery Data

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This study develops a theoretical model and then, using Canadian joint replacement surgery data, empirically tests the relationship between government policies that promote privately funded health care and patients' waiting time in the public health care system. Two policies are tested: one policy allows opt-out physicians to extra-bill private patients, and the other provides public subsidies to private patients. We find that both policies are associated with shorter public waiting time, and that the subsidy policy appears to be more effective in waiting time reduction than the extra-billing policy. Our findings are consistent with a dominant demand-side effect in that these policies would provide patients an option, and some incentive, to opt out of the public health system, shifting the demand from the public health system to the private care market.

Key words: health care waiting time; elective surgery; health policy; privately funded health service
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1. Introduction

Long waiting times in the public health system have been a source of public concern in many OECD (Organization for Economic Co-operation and Development) countries (e.g., Siciliani and Hurst 2005), and therefore a target for policy initiatives. In Canada, long waiting times for elective surgeries have led to several famous lawsuits and nationwide health reform debates.[1] In 2005, physician Jacques Chaoulli, together with joint replacement patient George Zeliotis, sued the Quebec government for banning patients' right to purchase private insurance to cover their private care expenses when they cannot obtain timely care in Canada's public health system. The Supreme Court's ruling forced the Quebec government to change its policies toward public waiting time and private care financing, which turned out to be highly contentious (Quesnel–Vallee et al. 2006). Due to its potential conflict with the Canada Health Act, some people believed that this ruling might lead to the dismantling of Medicare, Canada's publicly funded universal health insurance system, while others suggested that this was a strong call for reforming Medicare. Today, long waiting times and privatization are among the most controversial topics in the debates of health care reform in Canada and other OECD countries. In the United States, for example, Canada's waiting time problem has often been portrayed as a seemingly unavoidable consequence of universal health care.[2] In particular, the waiting time problem was used as evidence to counter the idea of introducing a public, government-administered health insurance to compete with private insurance (Blendon and Benson 2009).

Canadian federal and provincial governments have taken a number of measures to tackle the long waiting times in the public health system. Nonetheless, some people believe that the waiting time problem roots in Canada's single-payer health system, that is, a system with public health services being solely funded by the universal health insurance.[3] Although the Canada Health Act does not directly prohibit privately funded health services, many policies are in place to deter physicians from providing private care and restrain patients from purchasing private care. First, Canadian physicians are considered as for-profit entrepreneurs, so all health services are delivered privately. However, physicians need to choose between opt-in status and opt-out status. Opt-in physicians are entitled to bill the universal health insurance and are banned from receiving any private payments (i.e., out-of-pocket payments, or private insurance if permitted). In contrast, opt-out physicians are not allowed to bill the universal health insurance but can accept private payments. Second, many provinces do not allow opt-out physicians to extra-bill private patients, that is, charging patients an amount above what is paid in the public health system. Third, the ban on the purchase of private insurance also restrains patients' decisions to choose private care (Flood and Archibald 2001). Lastly, most provinces do not provide public subsidy to private care. As a result, patients literally do not have much of a choice in the face of long public waiting times.

Would broadening patients' access to private care be a remedy for the waiting time problem?[4] How would the availability and affordability of private care affect patients' waiting time in the public health system? The main objective of the present study is to assess the relationship between government policies that promote privately funded health care and the public waiting time. More specifically, we employ a unique data set from the Canadian Joint Replacement Registry (CJRR), Canada's only nationwide data registry that collects patient-level waiting time information for a specific surgical procedure. The data set encompasses joint replacement surgery records of nine Canadian provinces for 24 months.[5] The information contained in this data set has been used in the CJRR annual reports (e.g., CIHI 2008), but this is the first time that these data are analyzed with econometric methodologies. To generate the main hypothesis for empirical tests, we first develop a theoretical model of elective surgery in a public health system in which policies that induce privately funded health care affect both the demand for, and the supply of, public elective surgery. The waiting time, as an outcome of the demand-supply equilibrium, is then affected by the policies through the competing demand-side and supply-side effects: introduction of a private care sector shifts part of the demand for public care to the private market, thus mitigating congestion in the public health system. The private sector can, on the other hand, take resources away from the public system, which tends to reduce the public care supply and increase the public waiting time.

Two policies are then empirically tested: (i) allowing opt-out physicians to extra-bill private patients, and (ii) providing public subsidies to patients seeking private care. According to Flood and Archibald (2001), in provinces with the second policy in place, the government subsidizes any patient that seeks any type of publicly insured health service in the private sector up to the amount of public fee schedule. We find that both policies are associated with shorter public waiting times, and that the subsidy policy appears to be more effective in waiting time reduction than the extra-billing policy. Our finding is consistent with a dominant demand-side effect in the theoretical model in that these policies would provide patients an option, and some incentive, to opt out of the public health system, shifting the demand from the public health system to the private care market, while they did not appear to reduce the provision of public care very much.

These results shed some lights on Canada's health care reform debates. As indicated above, the two issues addressed in the paper, namely, public health care waiting time and private care financing, are at the heart of the health reform debates in Canada. However, no formal survey or study, in effect, has been conducted to investigate the status quo of private sector in Canada and the extent of private clinics' breach of the Canada Health Act. More generally, rigorous empirical work of Canadian evidence on the effect of private care financing on public waiting time is rare (Tuohy et al. 2004), which is due partly to lack of good data and partly to the complicated nature of health waiting times. In this regard, our study intends to fill the void.

In addition, this study contributes to the existing literature on private care financing in several ways. First, the study provides a more in-depth analysis compared to those studies that are based on case studies and simple data comparisons. Second, physician dual practice is prohibited in Canada, which is in contrast to the general practice in the other OECD countries. Therefore, the concern about the physician's incentive in manipulating patient waiting lists is minimal with our study.[6] Third, this study differs from the empirical studies that are based on survey data that are aggregated over different specialities, and from the empirical studies that test a different hypothesis, that is, whether long public waiting list is a driver to the demand for private health insurance (PHI). (A more detailed discussion on the existing literature is provided in section 2.) Lastly, the use of econometric techniques distinguishes our study from other studies that utilize Canadian data. For example, Cipriano et al. (2007) used Ontario joint replacement data to develop simulation models for waiting time prediction and policy evaluation.

The study is organized as follows. Section 2 briefly reviews the relevant literature on private care financing and discusses the contributions of the present paper. Section 3 develops a theoretical model of elective surgery in a public health system. Section 4 derives, based on the theoretical model, main hypotheses for empirical tests, develops the econometric specification, and discusses the data and variable construction. Section 5 presents the empirical results, and section 6 contains concluding remarks.

2. Literature Review

The existing literature on private health service is in general concerned with the financing of private care or the ownership of care provision. The present study focuses on the former aspect. Within the literature on private care financing, this study distinguishes from the previous empirical studies in the following ways. First, our study distinguishes from the literature that is based on case studies and simple data comparisons. By comparing country-level aggregated waiting list data, Tuohy et al. (2004) showed that jurisdictions with stronger private sector and private care financing, for example, England and New Zealand, appear to have longer waiting lists than those with weaker private sector and private financing, for example, Canada and the Netherlands. The authors also showed, nevertheless, that Australia, in which the parallel private hospital system is publicly subsidized, has a level of waiting lists similar to Canada. They suggested that the impact of private care financing on public waiting times rests on a complex political economy of waiting lists that requires much more research. In a similar spirit, Siciliani and Hurst (2005) used Australian data to show that the public waiting times for several procedures fall significantly following the Australian government's introduction of tax rebates to PHI in 1997–2000, which triggered an increase in the privately-funded share of activity in the years of 1999–2001. However, subsequent studies (Powers et al. 2003, Richardson and Segal 2004) questioned the substantiality of this waiting time reduction because PHI seems to induce a net increase in the total demand and the demand shifted from public hospitals to private hospitals are mostly same-day treatments instead of multi-day treatments. In summary, international evidences provide mixed findings on the association between private care financing and public waiting times, which call for more empirical analysis based on different contextual settings.

In addition to these mixed findings, it should be noted that the above studies use simple comparisons of aggregate data without controlling for, for example, demographic characteristics and contextual variables. By contrast, the control groups in our study are more homogenous with respect to these factors, owing to our unique institutional setting: our study uses data from a single country (Canada). Further, all Canadian provinces need to meet the principles of the Canada Health Act,[7] but provinces do regulate private care financing and physician's private practice in different ways (see Flood and Archibald 2001 for a review of these regulatory differences). The institutional setting thus presents us with an excellent research opportunity to investigate the effect of private care financing on public waiting time.

Second, our study also distinguishes from those studies that employ econometric techniques to investigate the association between private care financing and public waiting times, for example, Besley et al. (1998, 1999), Costa and Garcia (2003), Jofre-Bonet (2000), Schwierz et al. (2011), and Auteri and Maruotti (2012). All of these studies used survey or experimental data from European countries. Among them, the most relevant to our study appears to be Besley et al. (1998), which used British Social Attitudes Survey data to test the effects of PHI on the resource allocation in the public health system. In that study, long-term waiting lists were assumed as a proxy to health care resource allocation at the regional level. Its empirical results suggested that those regions with higher percentages of PHI have longer waiting lists, that is, less health care resource. Note that their waiting lists data are aggregated over all specialties, including those specialities that were not insured by PHI. Owing to this, Besley et al. (1998) did not test a more interesting hypothesis that longer waiting lists occur in the specialties insured by PHI. On this aspect, our study complements Besley et al. (1998) by providing an analysis using data based on a single specialty.

Besley et al. (1999), by using the same data set as Besley et al. (1998), investigated individuals' purchase of PHI, and suggested that longer waiting lists for public health service are associated with greater purchases of PHI. Jofre-Bonet (2000) provided an empirical analysis of the 1993 National Health Survey of Spanish data and showed that a reduction in the waiting time lowers the probability of buying PHI. In a similar spirit, Costa and Garcia (2003) tested a small set of survey data from Spain's Catalan health system and showed that patients' "perceived quality" gap between public and private care drives the demand for PHI.[8] These studies took a perspective that is different from ours. All these studies implicitly or explicitly considered the long public waiting list as a driver to the demand for supplemental PHI. In contrast, our perspective is that allowing private care financing may affect public waiting time.

Furthermore, Schwierz et al. (2011) addressed a related but different problem, that is, the prioritization of "public" and "private" patients, based on the institutional context of Germany. The authors tested experimental data and showed that patients paid by PHI are offered shorter acute care appointment times than patients paid by social insurance, which may be driven by the higher expected profitability of PHI holders relative to non-PHI holders for hospitals. Auteri and Maruotti (2012) tested survey data from Italian National Health System and showed that people with PHI have a higher probability of waiting and shorter waiting times. However, we consider their argument of a moral hazard effect for the higher probability of waiting not very convincing.[9] For these two studies, the waiting time comparison is between patients with and without PHI in a public health system. That is, in their settings, private care financing was already allowed, and the focus was then to look at different groups of people in the public health system. In comparison, our focus is to look at what would happen to the public health system if a certain extent of private care financing is allowed.

3. Theoretical Model

3.1. Demand-Supply Equilibrium

We first develop a demand and supply equilibrium for the surgery waiting time in the public health system. Let λ be the demand (patient arrival rate) for surgeries in the public health system. In general, λ depends on, among others, the price of public surgery, its service quality—which is to be approximated by its waiting time, denoted by w—and the availability and affordability of its substitutes. Since the price of public surgery remains constant (free of charge), the demand function can be expressed simply as:

$$\lambda = D(w, \eta; \Phi), \tag{1}$$

where η denotes the extent of privately funded care —that is, availability and affordability (lower price) of privately funded surgery, which serves as a substitute to public surgery—and Φ is the set of other factors that shift the demand. It is clear that $\partial D/\partial w < 0$, that is, when the public waiting time rises, less patients are willing to join the public system. Further, $\partial D/\partial \eta < 0$, that is, when the privately funded surgery becomes more available and affordable, less patients will stay in the public system.

Similar to Martin and Smith (1999), the supply of public surgery is modeled through the decision problem of the public health planner (e.g., the provincial health ministry). The planner's objective is to maximize the health surplus by allocating resources between public surgery and non-surgical activities, given demand for public surgery λ:[10]

$$\max_{\mu,\tau} U(\mu, \tau) = \lambda[R - \theta w] + v(\tau), \tag{2}$$

$$\text{s.t. } w = f(\lambda, \mu), \tag{3}$$

$$\mu + \tau \leq B_{HR}(\eta), \tag{4}$$

$$\mu + \tau \leq B_{NHR}(\eta), \tag{5}$$

where decision variables μ and τ are the amounts of supply for public surgery and non-surgical activities, respectively. R is the (gross) health benefit that a patient derives from public surgery, and θ is the cost of (per unit) waiting time. Thus, $\lambda[R-\theta w]$ is the net health surplus generated from public surgery. Further, $v(\tau)$ denotes the health surplus from non-surgical activities, with $\partial v/\partial \tau > 0$ and $\partial^2 v/\partial \tau^2 \leq 0$. Constraints (4) and (5) indicate that the supply of public health services is constrained by two types of resource, namely, human resource (B_{HR}) and non-human resource (B_{NHR}).[11] The former includes physicians, nurses, and other hospital personnel working in the public health system, whereas the latter refers to physical assets such as hospital facilities. While in general B_{HR} and B_{NHR} are functions of η, it is useful to consider that $\partial B_{HR}/\partial \eta \leq 0$ (i.e., if the availability/affordability of privately funded surgery increases, some human resources are likely to be attracted to the expanded private sector) while $\partial B_{NHR}/\partial \eta = 0$ (i.e., the availability/affordability of privately funded surgery has no impact on non-human resource). This simplified treatment highlights a difference in supply rigidity of the two resource types: the private sector is more easily to invest in physical assets using payments from private patients than in, for example, physicians. Finally, relationship Equation (3) follows queueing theories that waiting time is a function of the arrival (demand) rate and the service (supply) rate, with $\partial f/\partial \lambda > 0$, $\partial f/\partial \mu < 0$, $\partial^2 f/\partial \mu \partial \lambda \leq 0$ and $\partial^2 f/\partial \mu^2 \geq 0$.[12]

Solving the optimization problem proceeds as follows. Since the objective function (2) increases in μ and/or τ, one of the constraints (4) and (5) must be binding in optimality. Letting $B(\eta) \equiv \min\{B_{HR}(\eta), B_{NHR}(\eta)\}$, it must then follow that $\mu + \tau = B(\eta)$. Substituting $\tau = B(\eta) - \mu$ and Equation (3) into the objective function and taking the first- and second-order derivatives with respect to μ give rise to:

$$\frac{\partial U}{\partial \mu} = -\theta\lambda\frac{\partial f}{\partial \mu} - \frac{\partial v}{\partial \tau}, \tag{6}$$

$$\frac{\partial^2 U}{\partial \mu^2} = -\theta\lambda\frac{\partial f^2}{\partial \mu^2} + \frac{\partial v^2}{\partial \tau^2}. \tag{7}$$

According to the functional properties of f and v, it immediately follows that $\partial^2 U/\partial\mu^2 \le 0$ and thus the first-order condition $\partial U/\partial\mu = 0$ has a unique solution, which is to be denoted simply as μ. Such μ represents the (optimal) supply of public surgery and may be expressed as:

$$\mu = S(\lambda, B(\eta); \Psi), \tag{8}$$

where Ψ is the set of factors (other than λ and η) that shift the supply (e.g., R and θ).

It is of interest to examine the shape of supply function (8) with respect to λ and η. Substituting $\mu = S(\lambda, B(\eta);\Psi)$ into the first-order condition $\partial U/\partial\mu = 0$, totally differentiating the resulting identity and solving for $\partial S/\partial\lambda$ and $\partial S/\partial\eta$ yield:

$$\frac{\partial S}{\partial \lambda} = -\left(\theta\lambda\frac{\partial^2 f}{\partial\mu\partial\lambda} + \theta\frac{\partial f}{\partial \mu}\right)\Big/\left(\theta\lambda\frac{\partial^2 f}{\partial\mu^2} - \frac{\partial^2 v}{\partial\tau^2}\right) > 0, \tag{9}$$

$$\frac{\partial S}{\partial \eta} = -\frac{\partial^2 v}{\partial\tau^2}\frac{\partial B}{\partial\eta}\Big/\left(\theta\lambda\frac{\partial^2 f}{\partial\mu^2} - \frac{\partial^2 v}{\partial\tau^2}\right) \le 0. \tag{10}$$

The signs in Equations (9) and (10) can be readily verified by the functional properties of $f(\lambda,\mu)$,$v(\tau)$ and $B(\eta)$. Inequality Equation (9) shows that when the demand for public surgery rises, the health planner has an incentive to allocate more resources to the production of public surgery. This property is consistent with empirical evidence. For example, Windmeijer et al. (2005) used record level data from a Scotland hospital to estimate the responses of the supply and demand for secondary care to waiting lists size and waiting time. The authors found that increase in the waiting time and waiting list leads to increase in the supply. Perhaps the less obvious (and more interesting) result is the sign of $\partial S/\partial\eta$ in Equation (10). Here, there are two cases: (i) $\partial S/\partial\eta < 0$ if $B(\eta) = B_{HR}(\eta)$, that is, when an expanding private sector draws human resource away from the public health system, the supply of public surgery will be reduced; (ii) if $B(\eta) = B_{NHR}(\eta)$, however, then $\partial B/\partial\eta = \partial B_{NHR}/\partial\eta = 0$ and so $\partial S/\partial\eta = 0$. In this case, the existence of a private sector would not affect the supply of public surgery.

We assume the existence of the equilibrium in the same way as Linsay and Feigenbaum (1984), so the demand-supply equilibrium should simultaneously satisfy Equations (1), (3), and (8). Substituting Equation (1) and (8) into Equation (3) then yields the equilibrium waiting time as:

$$w = f(D(w, \eta; \Phi), S(D(w, \eta; \Phi), \eta; \Psi)). \tag{11}$$

3.2. Effects of Allowing Privately Funded Health Care on Public Waiting Time

Equation (11) implicitly determines the equilibrium waiting time, w, as a function of η:

$$w = h(\eta; \Phi, \Psi). \tag{12}$$

The comparative-static effect of η on w can be derived by substituting Equation (12) into Equation (11), totally differentiating the resulting identity and then reshuffling the terms, as:

$$\frac{dw}{d\eta} = \left\{\frac{\partial f}{\partial\lambda}\frac{\partial D}{\partial\eta} + \frac{\partial f}{\partial\mu}\left(\frac{\partial S}{\partial\lambda}\frac{\partial D}{\partial\eta} + \frac{\partial S}{\partial\eta}\right)\right\} \Big/\left(1 - \left(\frac{\partial f}{\partial\lambda} + \frac{\partial f}{\partial\mu}\frac{\partial S}{\partial\lambda}\right)\frac{\partial D}{\partial w}\right). \tag{13}$$

Note that the numerator of the right-hand side of Equation (13) consists of two competing effects:

$$\frac{\partial f}{\partial\lambda}\frac{\partial D}{\partial\eta} < 0, \tag{14}$$

$$\frac{\partial f}{\partial\mu}\left(\frac{\partial S}{\partial\lambda}\frac{\partial D}{\partial\eta} + \frac{\partial S}{\partial\eta}\right) > 0. \tag{15}$$

We call Equation (14) the "demand-side effect" and Equation (15) the "supply-side effect." A negative demand-side effect means that a greater extent of privately funded care induces patients to switch from public surgery to private surgery which reduces public waiting time. A positive supply-side effect implies, on the other hand, that a greater extent of privately funded care increases public waiting time. Two sources for this positive effect correspond to the two terms on the left-hand side of Equation (15): first, as the demand for public surgery falls when η rises, the health ministry allocates less resources to the production of public surgery, which in turn will increase public waiting time; second, when human resource is the bottleneck of the public health system, the increase of privately funded care "crowds out" the human resources used for the production of public surgery, which in turn will increase the public waiting time. This second effect will, however, disappear when non-human resource is the bottleneck of the public health system (recall in this case, $\partial S/\partial\eta = 0$).

If the demand-side effect dominates the supply-side effect, we must have

$$\begin{aligned}&\frac{\partial f}{\partial \lambda}\frac{\partial D}{\partial \eta}+\frac{\partial f}{\partial \mu}\left(\frac{\partial S}{\partial \lambda}\frac{\partial D}{\partial \eta}+\frac{\partial S}{\partial \eta}\right)<0\\ &\Rightarrow \frac{\partial f}{\partial \lambda}+\frac{\partial f}{\partial \mu}\frac{\partial S}{\partial \lambda}>-\frac{\partial f}{\partial \mu}\frac{\partial S}{\partial \eta}\Big/\frac{\partial D}{\partial \eta}\geq 0\end{aligned}. \quad (16)$$

Since $\partial D/\partial w < 0$, the denominator of Equation (13) is positive and it immediately follows that

$$\frac{dw}{d\eta}<0.$$

On the other hand, if the supply-side effect dominates the demand-side effect, we have

$$\begin{aligned}&\frac{\partial f}{\partial \lambda}\frac{\partial D}{\partial \eta}+\frac{\partial f}{\partial \mu}\left(\frac{\partial S}{\partial \lambda}\frac{\partial D}{\partial \eta}+\frac{\partial S}{\partial \eta}\right)\geq 0\\ &\Rightarrow \frac{\partial f}{\partial \lambda}+\frac{\partial f}{\partial \mu}\frac{\partial S}{\partial \lambda}\leq-\frac{\partial f}{\partial \mu}\frac{\partial S}{\partial \eta}\Big/\frac{\partial D}{\partial \eta}\end{aligned}.$$

In this case, the sign of $(\partial f/\partial\lambda+\partial f/\partial\mu\cdot\partial S/\partial\lambda)$ is undetermined, which makes the sign of the denominator of Equation (13) and further the sign of $dw/d\eta$ undetermined.

The effect of allowing privately funded health care on public waiting time depends, therefore, on the relative strengths of the competing demand-side and supply-side effects. The negative demand-side effect corresponds to the well-known supportive argument for private care—the existence of private care lessens the burden on the public health system. For example, Cullis and Jones (1985) argued, in discussing the competing explanations of public waiting list, that providing subsidy to private care is more efficient in reducing the public waiting time than increasing the budget of public care.

Iversen (1997) found that the net effect of a private sector on the public waiting time depends on the relative strengths of the negative *capacity effect* (the demand-side effect in this study) and the positive *demand side effect* (the supply-side effect in this study). The more elastic the demand for public care with respect to the public waiting time, the more likely the *demand side effect* is dominant, and thus the more likely the presence of a private sector will increase the public waiting time. The reason is that in Iversen (1997), the public health planner is assumed to behave like an economic agent who determines the public health expenditure by trading off the marginal benefit against the marginal cost of providing public care. The more elastic the demand is, the more likely the public planner would save costs, in the presence of a private sector, by lowering the public health expenditure to preserve a long public waiting time so that those delay-sensitive patients are forced to choose private care. In contrast, the model setup in this study is different. We assume that the public health planner aims to maximize patients' total benefits, as in the general case, so she has an incentive to maintain a public waiting time as short as possible. This objective is undermined when the private sector draws human resources from the public health system, a widely used argument against private care in the health reform debates (Tuohy et al. 2004). As a consequence, the elasticity of demand with respect to public waiting time does not play a major role in determining the sign of $dw/d\eta$.

4. Empirical Analysis

4.1. Policies and Main Hypothesis

Section 3 has shown that if the demand-side effect dominates the supply-side effect, an increase in the extent of privately funded care leads to shorter public waiting time. The impact is, on the other hand, undetermined if the supply-side effect dominates. Using Australian data, Siciliani and Hurst (2005, p. 211) showed that providing tax incentives to the purchase of supplemental PHI results in a dominant demand-side effect and reduces public waiting times. They suggested that governments should provide tax incentives for patients to purchase PHI and encourage patients to substitute private care for public care. For our empirical analysis, two government policies are used to proxy variable η discussed in section 3. Under the first policy, the government allows opt-out physicians to extra-bill private patients; as a consequence, more health service providers will be willing to participate in the private care sector. Under the second policy, the government provides public subsidy to patients seeking private care, reducing the price of privately funded surgery, and expanding the extent of privately funded care. In accordance with the empirical evidence and theoretical findings discussed above, the following main hypothesis is formed for our empirical tests:

HYPOTHESIS 1. *Both the policy of allowing opt-out physicians to extra-bill private patients and the policy of providing public subsidy to private patients are associated with shorter public waiting times.*

While both policies make privately funded surgeries more available, thereby shifting some demand away from the public health system, their impacts on the reduction in public waiting time may differ. While the subsidy policy improves both the availability and affordability of privately funded surgery, the extra-billing policy may reduce the patients' affordability (accessibility). As a result, we expect that the association identified in Hypothesis 1 would be stronger for the subsidy policy than for the extra-billing policy. It would be interesting to see if this difference exits in our empirical results.

4.2. Econometric Specification

To empirically test Hypothesis 1, function (12) is linearized, leading to the following econometric specification:

$$\begin{aligned} w_{it} = &\beta_0 + \beta_1 Extra\_billing_i + \beta_2 Subsidy_i \\ &+ \beta_3 Num\_surgeon_{it} + \beta_4 Hip_i \\ &+ \beta_5 Distance_{it} + \beta_6 BMI_{it} + \beta_7 Female_{it} \\ &+ \beta_8 Femaleage_{it} + \beta_9 Maleage_{it} \\ &+ \sum_{j=0}^{k} \beta_{10,j} N_{i,t-j} + \beta_{11} Month_t + \varepsilon_{it} \end{aligned}, \qquad (17)$$

where subscript i indicates province-joint, subscript t indicates month, binary variables *Extra_billing* and *Subsidy* denote the two government policies that, as mentioned above, proxy η, β's are the coefficients to be estimated, and ε is the error term. In addition to the two main explanatory policy variables, it behooves us to control for other potential drivers of waiting time w in Equation (17) in order to yield more robust causal inferences. These drivers (*control variables*) are now discussed:

(i) Number of surgeons: the number of orthopedic surgeons performing public hip/knee surgeries is used as one control variable for the supply of public care surgery. Other things being equal, the more surgeons working in a public health system, the shorter the public waiting times.

(ii) Hip/knee: CIHI (2008) found that, in general, hip patients have shorter waiting times than knee patients. However, the CIHI report does not elaborate on the causes of this difference.

(iii) Distance: our preliminary data analysis show that patients in provinces of Manitoba, Newfoundland and Labrador, Nova Scotia and Saskatchewan tend to have longer distances between home and hospital than patients in other provinces. Since these provinces have low population densities and lack of major urban areas where health care facilities are located close to the majority of population, their patients may have more difficulties in accessing to public health service than patients in other provinces.

(iv) Body mass index (BMI): Karlson et al. (2003) found that higher BMI significantly increases the risk of osteoarthritis. Obese patients may be requested to lose weight prior to undergoing surgeries to improve the outcome of the treatment (CIHI 2008) and thus these patients may experience longer waiting times. In our data set, however, the BMI information is missing in a large number of records for some provinces (the percentage of surgeries with BMI information ranges from 39% for Ontario hip surgery to 91% for Saskatchewan knee surgery). The available data show that the distributions of BMI have small variations. For example, the coefficient of variation ranges from 0.18 for Alberta knee patients to 0.27 for Manitoba hip patients. Furthermore, the average BMIs across provinces are quite close: the average BMI ranges from 28.03 for British Columbia hip patients to 32.91 for Saskatchewan knee patients. Therefore, we assume that the available BMI data are representative, so the average BMIs can be calculated based on these data.[13] Note that the *BMI* variable in regression (17) is both province-joint and time dependent.

(v) Age and gender: CIHI (2008) reported waiting time differences by gender and age, and age difference by joint type. For example, CIHI (2008) found that female patients tend to have shorter waiting times than male patients, and that knee patients are significantly older than hip patients. Karlson et al. (2003) found that older age significantly increases the risk of osteoarthritis, the most commonly reported diagnosis of joint replacement patients. In Equation (17), we use variables *Femaleage* and *Maleage* to isolate the age effect by gender, and use the *Female* variable (the percentage of female patients in total patients) to capture the gender effect.[14]

(vi) Count of incoming patients: as will be discussed below, the time-series variations of prospective waiting time depend on the history of arrivals and services. If each surgeon is considered as a separate server, then the basic queueing theories imply that arrivals have positive lagging effects on (prospective) waiting times. We hypothesize that such positive lagging effects also prevail in a public health system with many surgeons in aggregation. Hence, we propose lag variables N_{t-j}, $j = 0, 1, 2, \ldots, k$ in the econometric specification. We run regressions with different $k's$, the number of lags, to test the lagging effect, and check the robustness. Similar to the construct of prospective waiting time discussed below, the construct of the number of arrivals is also subject to the problem of data truncation. To remedy this problem, we estimate the number of arrivals prior to April 2005 based on data inclusion. The estimation of the number of arrivals prior to April 2005 works as follows: for month t

prior to April 2005, we estimate that $y\%$ of patient records are not included in our data set. Let c be the count of surgery records for month t and thus we estimate $N_t = c/(1-y\%)$. We normalize the number of arrivals by province-specific population and data inclusion. Nevertheless, Table S1 of Appendix S1 shows that many provinces have a mild increase in data inclusion from 2005/6 to 2006/7, which may counteract the data truncation effect.

(vii) Trending: our preliminary analysis of waiting time data shows a slight declining trend over time for some provinces, for example, Alberta, British Columbia, and Manitoba. Provincial governments' short-term measures to deal with patient backlog may contribute to this time trend (and so trending may serve as a supply-side control variable). On the other hand, the data truncation problem may also contribute to the declining trend of waiting times and patient arrivals. When both the dependent and explanatory variables are trending, Wooldridge (2009) suggested adding a time trend variable to obtain a de-trending interpretation of the regression. Therefore, we add a time trend variable, *Month*, which measures the number of months after March 2005, to the econometric specification Equation (17).[15]

Except for the interaction effects of age and gender, no other interaction effects have been suggested by the existing literature or found statistically significant in our preliminary analysis of differences between means.

4.3. Data and Variable Construction

The data for empirical tests are comprised of 29,369 joint replacement surgery records of nine Canadian provinces based on admission date from 1 April 2005 to 31 March 2007. Each record contains a patient's demographic information (age, gender, BMI, home province, distance from home to hospital), surgical information (hip/knee, primary surgery/revision, surgeon ID),[16] and waiting time information (the calendar month of making decision for surgery, the calendar month of admission to hospital,[17] and waiting time defined as the number of days from decision date to admission date). Only primary surgery data are included for our analysis as revision surgeries are to repair the primary joint replacements. As a consequence, a total of 2277 revision records, together with 51 erroneous primary records which have decision dates after admission dates, are removed.

Canadian orthopedic surgeons submit their joint replacement surgery records to CJRR on a voluntary basis. Depending on a province's surgeon participation rate and data submission, the percentage of surgeries included in our data set (to be referred to as "data inclusion" hereafter) varies by province, ranging from 3.6% of hip/knee replacements of Ontario in 2005/6 to 70% of knee replacements of New Brunswick in 2006/7.[18] Due to the voluntary data submission, it behooves us to assess the representativeness of the data. First, we compared our waiting time data with the waiting time data published in Fraser Health Institute's (FHI) annual waiting time surveys of 2005 and 2006 (Esmail, and Walker 2005,, 2006). We found a strong consistency: the median waiting times (at province/year level; hip and knee combined) of our data are strongly and positively correlated with the median waiting times of arthroplasty surgeries (at province/year level; hip, knee, ankle, and shoulder surgeries combined) from FHI's surveys—the correlation is 0.78 for 2005 and 0.87 for 2006. To our best knowledge, FHI's annual surveys are the only source that contains waiting time data at the specialty level using a consistent method.

Next, we compared the patient demographics of our data, that is, age, gender, and joint type distributions of the 29,369 records, with the patient demographics of Hospital Morbidity Database (HMDB) and Discharge Abstract Database (DAD), both of which contain all joint replacement hospitalizations during the same period as our study period. The patient demographics of these two sources are very similar, if not closely matched. To deal with outliers of waiting time and distance data, we followed the common procedure of winsorizing the data—we replaced the top (bottom, respectively) 1% of the data by the 99 (1, respectively) percentile value for each province-joint pair. We also experimented with 2.5% and 5% cutoffs. The exact location of winsorization did not affect our results, however.

The dependent variable, w_{it}, can be measured in two fashions: *prospectively* and *retrospectively*. The *prospective* (*retrospective*, respectively) waiting time defines w_{it} as the average waiting time of patients who *arrive* (*depart*, respectively) in calendar month t. We decided to use the prospective waiting time for the following reasons. First, the prospective waiting time is consistent with our (implicit) assumption in the demand function that patients make their decisions based on the expected waiting time looking forward. Second, the prospective waiting times constructed at the province/joint/month level reflects the transient states of a queueing system (i.e., the public health system here). These transient waiting times have two components: one is the equilibrium part, that is, w in Equation (12), and the other is the time-series variations. Queueing theories suggest that the time-series variations depend on the history of arrivals and services. To account for these variations, we added the explanatory variable N in the econometric specification, as discussed above. This variable counts the number of patients *entering* the waiting list in month t. It is of our interest to see whether classic queueing results also prevail in health care data.

Nevertheless, the prospective waiting times come with the cost of data truncation. For example, the patient records with decision dates falling within the 24-month period but admission dates falling after March 2007 are not included in our estimation data as a result. Data truncation induces a bias toward the underestimation of w_t, especially when the estimation moves closer to March 2007. To ensure that the time-series data of different *cross section* (province-joint) units are comparable, for each cross-section unit we only include the data of those months in which at least 90% of the incoming patients are served by the end of the 24-month period. For instance, for cross-section unit *Alberta-hip*, 90% of its patient records have waiting times less than or equal to 11 months, that is, it is reasonable to argue that at least 90% of the patients arriving in the first 13 months of our study period were served by the end of the 24-month period. Therefore, the first 13 months of data (i.e., April 2005 to April 2006) are used for *Alberta-hip*. Table S2 in Appendix S1 shows the number of months included in the analysis for each cross-section unit, which ranges from 0 of *Saskatchewan-knee* to 18 of *Newfoundland-hip/knee*.

In econometric specification Equation (17), therefore, subscript i is the indicator of a province-joint pair and t is the indicator of calendar month, while all the variables, except for the two policy variables and the variable of the number of working surgeons, are calculated in the prospective manner. The two policy variables, *Extra_billing* and *Subsidy*, are two dummy variables indicating, respectively, the provinces that allow opt-out physicians to extra-bill private patients and the provinces that provide public subsidy to private patients (Flood and Archibald 2001). CJRR assigns a surgeon ID for each patient record in the data set, so the *Num_Surgeon* variable is the number of working orthopedic surgeons at the joint/province/month level. These numbers are further standardized by province population and adjusted for data inclusion.

Table 1 lists the name, definition, and source of the variables and summary statistics. Table 2 shows the mean waiting times by province and joint, and indicates

Table 1 Variables Available for Inclusion in Regression Models

Variable	Definition	Mean	Sd	Minimum	Maximum	Source
w	Mean waiting time (in days)	164.02	63.06	63.46	338.26	CJRR
Extra_billing	Dummy variable. 1 if opt-out physicians are allowed to extra-bill private patients and 0 otherwise.	0.719	0.45	0	1	Flood and Archibald (2001)
Subsidy	Dummy variable. 1 if public subsidy is provided to private patients and 0 otherwise.	0.18	0.38	0	1	Flood and Archibald (2001)
Num_Surgeon	Number of working surgeons per 10^5 population adjusted for data inclusion	2.03	0.56	0.60	3.60	CJRR
Hip	Dummy variable. 1 if hip surgery and 0 otherwise.	0.54	0.49	0	1	CJRR
Distance	Mean distance from home to hospital (in kilometers)	39.47	27.48	13.78	304.16	CJRR
BMI	Mean BMI of incoming patients	30.40	2.15	24.25	36.28	CJRR
Female	Percentage of female patients	0.58	0.09	0.28	0.93	CJRR
Maleage	Mean age of male patients	65.17	4.37	41.75	75.67	CJRR
Femaleage	Mean age of female patients	67.22	2.98	56.4	74.54	CJRR
N	Number of incoming patients per 10^5 population adjusted for data inclusion	8.59	2.94	2.70	17.50	CJRR
Month	Number of months after April 2005	6.69	4.14	1	18	CJRR

Table 2 Mean Waiting Times and Policies By Province

Province	AB	BC	MB	NB	NL	NS	ON	QC	SK
Mean waiting time—hip replacement (in days)	151	203	252	173	95	234	97	138	324
Mean waiting time—knee replacement (in days)	226	255	307	212	108	271	104	165	442
Opt-out physicians allowed to extra-bill private patients?	Y	Y	N	Y	Y	N	N	Y	Y
Public subsidy provided to private patients?	N	N	N	N	Y	N	N	N	N

AB, Alberta; BC, British Columbia; MB, Manitoba; NB, New Brunswick; NL, Newfoundland and Labrador; NS, Nova Scotia; ON, Ontario; QC, Quebec; SK, Saskatchewan; Y, yes; N, no.

whether a province allows opt-out physicians to extra-bill private patients and/or provides public subsidy to private patients. In consistent with the discussions above, we hypothesize that arrivals have positive lagging effects on the public waiting time, and older patients, male patients, knee patients, and patients having longer home-to-hospital distances have longer waiting times.

To properly employ our data and to check the robustness of the analysis, several econometric issues need to be addressed. First, due to the non-variation of the two policy variables, we are unable to fully employ panel data econometric techniques, that is, the fixed-effects model. We, therefore, resort to the feasible generalized linear square (FGLS) model.[19] Second, the FGLS model uses panel-specific AR(1) autocorrelation structure and allows for heteroskedasticity, but assumes no cross-sectional correlation. To account for the unobserved cross-section specific effects, we also examine our data using the random-effects specification, which clusters standard errors at the province-joint level. Third, the inclusion of the *Num_Surgeon* variable might introduce a potential endogeneity problem. Unfortunately, we are unable to find effective instrumental variables to remedy this problem, owing to data limitation. Hence, we only report regression results without this variable being instrumented. Fourth, due to the general multicollinearity problem between lag variables, the *t*-statistics of lag variables are often less significant. Wooldridge (2009) suggested using the Wald test to test the joint significance of lag variables. The Wald test is also used to test the joint significance of variables *BMI* and *Maleage*, which are found to be significantly positively correlated. Finally, we have considered other factors that may potentially affect public waiting time, especially those that may influence the supply side of surgery, for example, the public health expenditure and the ratio of general physicians to orthopedic surgeons. However, data for these variables are very limited and only available at the aggregated level, so we decided not to report the results with these variables being included in the regressions. More discussions on these variables and their test results are provided in Appendix S1.

5. Results and Discussion

Table 3 shows the results of both the FGLS models and the random-effects models. For both specifications, results of the Wald test suggest that only when the 3-month and 4-month lags of arrivals are included, the coefficients of lag variables are jointly significant.[20] The estimated coefficients of arrivals and lags are mostly positive, indicating positive lagging effects of arrivals on the public waiting time. Thus, it seems that classic queueing results also prevail in our health care data. Furthermore, the results suggest that if there is one more patient arrival in the current month, patients arriving 3 or 4 months later should expect an extra delay of 1–3 days over the average waiting time.

The estimated results of other explanatory variables are as follows. The coefficients of *Hip* are negative for all the specifications but are only statistically significantly different from zero for the FGLS model with 3-month lags of arrivals. The results suggest that knee patients generally have to wait one to three weeks longer than hip patients. This finding is consistent with the CJRR annual report (CIHI 2008). The coefficients of *Distance* are positive for all the specifications but are only statistically significant for the FGLS models, suggesting that distance affects patients' waiting time mainly through the province-specific mean, a finding that is consistent with our preliminary analysis of distance data.

The signs of *BMI* coefficients vary across the specifications but are statistically insignificant. The coefficients of *Female* are negative but insignificant for all the specifications. The negative sign is consistent with the findings of CIHI (2008)—female patients tend to have shorter waiting times than their male counterparts. This may be due to the higher obesities of male patients, as the *BMI* and *Maleage* variables are significantly positively correlated. As obesity is positively correlated with poor health condition, obese patients may require more medical treatments prior to surgeries. However, the *BMI* and *Maleage* variables are not jointly significant. Therefore, the relations between obesity, patient's age, and patient's waiting time deserve further investigation. The regression results, on the other hand, exhibit mixed age effects on waiting times. The coefficients of *Femaleage* are negative but mostly insignificant, while the coefficients of *Maleage* are inconclusive statistically. Finally, the coefficients of *Month* are negative and significant for all the specifications, indicating a declining time trend. The data truncation problem may contribute to this time trend. Perhaps more importantly, in September 2004, the federal and provincial governments set up the Wait Times Reduction Fund (in *the 10-Year Plan to Strengthen Health Care*) so as to build capacities to address waiting time problems. All the provincial governments were required to pledge additional funding to this initiative. Joint replacement surgery was identified as a target area for this fund. These government-led initiatives may affect the joint replacement surgery waiting times through the following channels: first, the fund could directly increase the provision of joint replacement surgeries; second, under the pressure of achieving meaningful waiting time reductions by March 2007, provincial governments

Table 3 Regression Results

	FGLS model		Random-effects model	
Intercept	324.0	306.7	186.8	180.5
	(70.10)***	(67.75)***	(79.88)*	(80.93)*
Extra_billing	−44.09	−40.00	−22.47	−20.00
	(7.874)***	(7.585)***	(34.22)	(34.51)
Subsidy	−83.20	−79.05	−80.51	−77.33
	(8.312)***	(8.123)***	(12.67)***	(12.05)***
Num_Surgeon	−2.554	−5.350	−5.220	−5.414
	(3.531)	(3.497)	(6.049)	(6.458)
Hip	−16.96	−9.988	−14.13	−9.957
	(7.420)*	(7.187)	(25.13)	(25.20)
Distance	0.601	0.513	0.247	0.284
	(0.123)***	(0.121)***	(0.161)	(0.181)
BMI	−1.030	−0.794	0.911	0.768
	(1.060)	(1.009)	(1.122)	(1.050)
Female	−16.86	−24.95	−23.80	−30.84
	(18.06)	(17.45)	(20.29)	(23.24)
Maleage	−0.284	−0.0696	0.166	0.188
	(0.488)	(0.484)	(0.536)	(0.619)
Femaleage	−1.312	−1.538	−0.145	−0.206
	(0.669)	(0.661)*	(0.821)	(0.841)
N	1.367	1.178	0.814	0.588
	(0.736)	(0.711)	(1.195)	(1.199)
N_{-1}	1.277	1.481	0.327	0.441
	(0.704)	(0.680)*	(1.178)	(1.192)
N_{-2}	0.452	0.500	−0.997	−0.675
	(0.698)	(0.686)	(1.199)	(1.329)
N_{-3}	3.550	3.542	2.547	2.472
	(0.699)***	(0.669)***	(0.891)**	(0.862)**
N_{-4}		2.405		1.506
		(0.695)***		(1.523)
Month	−2.486	−2.876	−1.858	−2.215
	(0.587)***	(0.581)***	(0.700)**	(0.749)**
Observations	184	184	184	184
χ^2	331.65	372.96		
R^2			0.38	0.39
Wald test that coefficients of *N* and its lag variables are all equal to zero (i.e., Prob $> \chi^2$)	0.000	0.000	0.014	0.013
Wald test that coefficients of *BMI* and *Maleage* are all equal to zero (i.e., Prob $> \chi^2$)	0.471	0.652	0.551	0.513

***, **, and * denote statistical significance at the 1%, 5%, and 10% confidence levels, respectively. Standard errors are shown in parentheses.

might have shifted some capacities used previously by other services to joint replacement surgery. Johnson et al. (2007) showed that between 2001/2 and 2005/6, there was a dramatic increase in the age-standardized rate for joint replacement surgery, but a decrease for other types of surgery.

Regarding our main explanatory variables, *Extra_billing* and *Subsidy*, the results of both the FGLS models and the Random-effects models suggest that the policies that induce private care financing are associated with shorter public waiting times, supporting Hypothesis 1. Furthermore, the results of *Subsidy* are more robust than those of *Extra_billing*, noting in particular that the magnitude of *Subsidy* coefficients is 2–4 times greater than that of *Extra_billing* coefficients, indicating that providing subsidy to private patients is associated with a greater waiting time reduction than allowing opt-out physicians to extra-bill private patients. This result is consistent with our expectation: as discussed earlier, the result may be explained in part by the demand-side effect in Equation (16). The policy of allowing opt-out physicians to extra-bill private patients only provides patients with an additional option of paying in full amount for private care (improving the availability of private care but not its affordability), while the policy of subsidizing private patients up to the amount of public fee schedule provides patients a more significant incentive to opt out of the public health system (improving both the availability and affordability of private care) with a larger amount of demand being induced from the public health system to the private care market as a result. For a given supply-side effect, therefore, the policy of subsidization would induce a stronger demand-side effect than the policy of allowing extra-billing.

The coefficients of *Num_Surgeon* are all negative but not statistically significant, a finding that is consistent with the common wisdom that a public health system with a larger number of working surgeons should supply more services, thereby leading to shorter patient waiting times. As mentioned earlier, there is a potential endogeneity problem between the number of surgeons working in the public health system and the private care financing policies because the existence of a private sector may attract surgeons from the public sector. Unfortunately, we were unable to find effective instrumental variables to mitigate this problem. Nonetheless, the endogeneity problem, if any, would bias the estimation toward a positive policy effect on waiting time. Therefore, although the endogeneity problem might be present, the signs of the policy variables are expected to be consistent and robust.

To test the robustness of our findings, we had also conducted various sensitivity analysis:

(I) To account for the likely high and uncontrolled clustering of dependent variables, we run regressions with standard errors clustered at the province-joint level. Compared to the results of Table 3, the only difference in these tests is that the coefficients of *Extra_billing* become insignificant in the FGLS models with clustering.

(II) Autocorrelation is a main concern when running regressions on time-series data. We applied the Wooldridge (2002, Pp. 282–283) test for autocorrelation to our data set, and the test result did not reject the null hypothesis of no first-order autocorrelation. To avoid the potential correlation of error terms between joint types of the same province, we further checked the problem by running the regressions solely on hip or knee data. The test results showed that the regressions using knee data yield more significant results than the regressions using hip data.

(III) In Table 3, we only presented the results of the regressions with a common time trend as we believe that a common time trend is better to reflect the macro movement of waiting times over the 24 months period. To account for province-specific waiting time trends, we replaced a common time trend by eighteen province-joint specific time trends in the regressions. The results showed a declining trend for five province-joint pairs, an increasing time trend for two province-joint pairs and no time trend for others.

In summary, the sensitivity analysis shows that the results presented in Table 3 are robust.

6. Concluding Remarks

How the availability and affordability of private health care affect patient waiting time in the public health care system has been extensively debated in Canada and other OECD countries. Our main objectives in writing this study are to assess the relationship between government policies that promote privately funded health care and the public waiting time, and to shed some light on the debates as well as contribute to the existing literature. A theoretical model was developed and its empirical implications were drawn. We then investigated the relationship using a sample of joint replacement surgery data from Canada. We found that both the provincial policy of allowing opt-out physicians to extra-bill private patients and the provincial policy of providing public subsidy to private patients are associated with shorter waiting times, and that the subsidy policy is more effective in reducing waiting time than the extra-billing policy. These findings are consistent with a dominant demand-side effect in that both policies would provide patients an opportunity to opt out of the public health system, shifting the demand from the public health system to the private care market, while they do not appear to reduce the provision of public care very much. The combined effects then lead to reduced public waiting times. The result of subsidization being a more effective policy is consistent with the conclusion drawn by Flood and Archibald (2001) that public funding to private care is the key in making private care an option for patients.

This study has also raised a number of issues and avenues for future research. First, if public subsidy to private care reduces the public waiting time, what should be the optimal amount of such subsidy? Should the subsidy be adjusted dynamically to achieve the maximum amount of health care surplus? These questions have been addressed, in part, in Chen et al. (2011). Second, the Fixed-effects model is not applicable in this study due to the non-variation of the policy variables in our data set; the waiting time data with a longer time horizon may, however, change this and allow one to employ the panel data methodology. Third, more data may also allow one to empirically isolate the demand-side effect from the supply-side effect. While our findings are consistent with a dominant demand-side effect, it does not necessarily imply that the demand-side effect is dominant, since we cannot exclude the possibility that the supply-side effect is dominant. (Recall from the theoretic model of section 2 that when the supply-side effect is dominant, the impact of introducing private care on waiting time is undetermined.) To separate the demand-side and supply-side effects would require the estimation of a structural-equations system, a task that necessitates more detailed data. We see these exercises as a natural extension of the analysis presented here, although beyond the scope of the present study.

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Notes

[1]These lawsuits include Chaoulli vs. Quebec in 2005 (joint replacement), Murray vs. Alberta in 2006 (joint replacement), and McCreith and Holmes vs. Ontario in 2007 (cancer).

[2]For instance, Wendell Goler, *Canada's Health System Informs U.S. Health Care Debate*, http://www.foxnews.com/politics/2009/08/12/ (12 August 2009).

[3]See, for example, Evans (2000) for an overview of the structure and funding of Canada's health system.

[4]Based on the institutional context introduced above, the difference between public care and private care rests predominantly on the financing side. In the remainder of the study, therefore, "private care" or "private health service" refers to services that are funded through private sources, while "public care" or "public health service" refers to services that are funded through the universal health insurance plan.

[5]Canada is composed of 10 provinces—the one province that is left out of the CJRR data is Prince Edward Island (PEI) and three territories (Northwest Territories, Nunavut, and Yukon). The combined population of PEI and the three territories accounts for only 0.8% of the Canadian national population.

[6]For instance, if a physician is allowed to work in both the public and private sectors, he or she might strategically manipulate the public waiting list so as to increase the demand for private care.

[7]The five principles are public administration, comprehensiveness, universality, portability, and accessibility. See Health Ministry of British Columbia, "Canada Health Act and Its Principles,"http://www.health.gov.bc. ca/library/publications/year/2007/conversation_on_health/PartII/PartII_HealthAct.pdf (accessed 10 April 2012).

[8]The perception of quality is defined as patients' overall assessment of certain observable service characteristics including, but not limited to, waiting time, flexibility to cope with specific individual needs and the amount of attention to "non-clinical" health care outcomes.

[9]According to the authors, "...... it becomes clear that owning a private insurance increases the probability of waiting and, thus, no benefits can be identified. This can be considered as a moral hazard effect: individuals with insurance coverage, and therefore lower opportunity costs, show a higher propensity to use the [public] health care services regardless of their health status. For this reason, they are likely to wait [in the public sector] ......" (Auteri and Maruotti 2012, P. 462). One might wonder, however, that since the Italian government does not charge patients for their use of public health service, those patients without PHI can abuse the public health care as well, that is, to use public health care service regardless of their health status.

[10]It is reasonable, in our setting, to assume that λ is given in the public health planner's decision making. For example, the health ministry does its planning based on the demand forecast for a relatively long-time horizon, for example, one year. Once the planning decisions are made, they are unlikely to be affected by short-term demand fluctuations.

[11]In addition, they imply that it costs one unit of human resource and one unit of non-human resource to supply one unit of public care (regardless of the public surgery or non-surgical activities). This simple production technology is adopted for simplicity of the analysis (without loss of its generality).

[12]These second-order derivative signs hold for general queueing models.

[13]For example, suppose that there are c surgeries for month t, of which the BMI information is only available for c_b surgeries. The value of the *BMI* variable for month t is the mean of these c_b surgeries.

[14]For example, suppose that there are c surgeries for month t, among which c_m is the number of male patients and c_f is the number of female patients, such that $c_m+c_f = c$. For month t, the value of the *Female* variable is the percentage of female patients, that is, c_f/c, the value of the *Femaleage* variable is the mean age of these c_f female patients, and the value of the *Maleage* variable is the mean age of these c_m male patients.

[15]We have also tested specifications using indicator variables for calendar months. The main results of this study remain unchanged however.

[16]Other surgical information is excluded from our analysis due to a large percentage of missing data.

[17]CJRR only agreed to provide decision dates and admission dates in the form of calendar month.

[18]The detailed information of data inclusion is shown in Table S1 of Appendix S1.

[19]For other studies that utilize panel methods to empirically investigate waiting time issues, see Martin and Smith (2003) and Siciliani and Martin (2007).
[20]We also used the *F*-test to test the specifications—comparing the log-likelihood values with and without the lag variables. Similar significance test results were obtained.

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Supporting Information

Additional Supporting Information may be found in the online version of this article:

Appendix S1: Variable Constructs and Consideration of Additional Variables.