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The demand for private health insurance: do waiting lists or waiting times matter?

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Abstract

Besley, Hall, and Preston (1999) estimated a model of the demand for private health insurance in Britain as a function of regional waiting lists and found that increases in the number of people waiting for more than 12 months (the long-term waiting list) increased the probability of insurance purchase. In the absence of waiting time data, the length of regional long-term waiting lists was used to capture the price-quality trade-off of public treatment. We revisit Besley *et al.*'s analysis using Australian data and test the use of waiting lists as a proxy for waiting time in models of insurance demand. Unlike Besley *et al*., we find that the long-term waiting list is not a significant determinant of the demand for insurance. However we find that long waiting times do significantly increase insurance. This suggests that the relationship between waiting times and waiting lists is not as straightforward as is commonly assumed.

Keywords: waiting time, waiting lists, health insurance, regional aggregation

1 Introduction

Admission via a waiting list is commonly used as a rationing device for non-emergency procedures in the healthcare sector where public services are free. In OECD countries where waiting lists are used, average waiting times for procedures often exceed six months (Siciliani and Hurst, 2003). Delays in medical treatment can prolong suffering, decrease earning capacity, and cause deterioration in quality of life. It is hypothesised that consumers buy private health insurance (PHI) covering private inpatient care, to give them the option to receive treatment as a private patient and avoid long delays associated with free treatment (Colombo and Tapay, 2004; Harmon and Nolan, 2001).[1] Private insurance smooths income by reducing large, unexpected, out-of-pocket costs at the point of purchase of private treatment. A seminal paper that supports this hypothesis is Besley, Hall, and Preston (1999).

BHP estimated a model of demand for PHI using a sample of British individuals over the period 1986–1991. Insurance demand is modelled as a function of individual characteristics and regional waiting list variables: the total inpatient waiting list and the long term list. The long-term waiting list was defined as the number of individuals (per thousand of the population) who had been on the waiting list for at least 12 months. On average the long term waiting list accounted for 20% of the total waiting list. Controlling for household income and other demographics and the size of the total waiting list, BHP predicted the insurance rate to increase by 2% per additional long-term patient (per thousand population). This finding has been interpreted as indicating that the demand for private treatment is responsive to the quality of public health care. The result has been used to support arguments for subsidising PHI in order to shift the health care burden to the private sector and reduce public hospital waiting times (Siciliani and Hurst, 2003, 2005). In fact, between 1997 and 2000, the Australian government embraced this strategy to reduce public hospitals' waiting times (Willcox et al., 2007; Duckett 2005). Other researchers have interpreted BHP's result as indicating a role of the public health system in producing better quality health care in the private sector (Bethencourt and Galasso, 2008).

In this paper, we use Australian data to replicate the BHP analysis and test the usefulness of waiting lists as a proxy for waiting time. We use administrative data

[1] Other benefits of private treatment may include choice of doctor and quality of accommodation during an inpatient stay.

on planned admissions to New South Wales (NSW) public hospitals during the period 2004-2005. This data contains information on patient-level waiting times and can be used to construct regional waiting time and waiting list data. Using information on residential postcode, regional waiting time and list variables are combined with individual data on insurance status and socio-demographic characteristics from the Household, Income and Labour Dynamics in Australia (HILDA) data set year 2005/2006. Our use of Area Health Services (AHS) for waiting list and waiting time variables corresponds to the use of regional health authorities by BHP. In the insurance models we include similar controls to those used by BHP in their model of individually-purchased health insurance demand.

It is potentially misleading to use waiting lists and waiting times interchangeably. Cullis and Jones (2000) state that "(m)any have examined numbers on lists rather than average time (or the distribution of time) spent on lists. But it is the latter factor which affects the behaviour of demanders or their agents" [p.1229]. Waiting time statistics based on treated patients have been found more accurate than those derived from waiting lists, where longer waiting time are likely to be over represented due to overlaps between census dates (Godden and Pollock, 2008; Dixon and Siciliani, 2009). The two measures can move in different directions (Siciliani, 2008), or be independent of each other (Newton et al., 1995). For example, in England, the NHS waiting list declined steadily (from more than 1.2 million in 1997 to 1 million in 2002 and to just 800,000 patients in 2004) while the average waiting times remained relatively stable (NHS, 2009; Appleby, 2005). Waiting times fell only when the British government introduced policies on maximum waiting times (Willcox et al., 2007). In contrast, in Australia additions to the waiting lists steadily increased from 584,000 in 2000 to over 740,000 in 2007/2008, but the median waiting times remained relatively stable at about one month (AIHW, 2000; 2004; 2006; 2008).[2]

Consistent with the BHP results, we find that the long-term waiting list is positively related to demand for private health insurance, but it is never statistically significant. In contrast, we find that long waiting times significantly increase insurance demand. Specifically, insurance demand is sensitive to the upper tail of the waiting time distribution, as measured by the 90th percentile waiting time. These findings have policy implications beyond the testing of a previous result. First, they show that waiting lists and waiting times do not necessarily measure the same phenomena. Second, our analysis suggests that the impact of waiting time on PHI demand is

[2] Hanning (2002) suggests that additions to the waiting list measure demand.

more modest than suggested by BHP. This result is consistent with previous studies which find a weak relationship between demand, supply, and waiting times (Martin and Smith, 1999, 2003; Goddard and Tavakoli, 1998). Johar *et al.* (2010) estimate the effect of individual waiting time and PHI demand and also find no effect of expected waiting time at the individual level. The lack of responsiveness of PHI demand to waiting times suggests that proposals to control waiting times through the private health insurance market are unlikely to be successful. Heterogenous motives of insurance purchase may be a factor that weakens the relationship between waiting time and insurance.

2 LITERATURE

There is only one previous study that uses waiting times, instead of waiting lists, to predict PHI demand. Jofre-Bonet (2000) use the Spanish National Health Survey and construct waiting time for respondents who were on the waiting list and were admitted during the study period. Individual waiting times are then aggregated to province-level. Consistent with BHP, the study found a large effect of waiting time on the insurance decision. She finds the probability of PHI purchase increases by 0.3-0.4 percentage points for an additional 15 days of waiting. This result however is significant only at the 10% level and is sensitive to model specification and sample selection. In particular, a highly significant result is found only when the respondent is not a household head (i.e., they are likely to be beneficiaries rather than the purchaser of the insurance). In addition, in-sample calculations of waiting times can be problematic if the hospitalised population is different from the general population; some illnesses will be over- or under-represented in the general population data.

King and Mossialos (2005) update the BHP analysis using 1997–1999 NHS data. To measure long wait, they used the proportion of patients waiting over 6 months for an inpatient stay. They find that long waiting time influences individually-purchased health insurance decisions but not employer-financed insurance decisions. They find that a 1% increase in long wait was associated with a 4% increase in the odds-ratio of insurance purchase.

Other studies, which do not have waiting list data, rely on measures of perceived quality of public hospitals (Costa-Font and Font-Vilalta, 2004; Costa and Garcia, 2003). Costa and Garcia (2003) use expressed satisfaction with the public system and find that perceived lower quality increases the probability of purchasing PHI in Catalonia. At the mean, public care quality was perceived to be 2 index points (on a scale of 1 to 10) lower than private care. Holding quality of public care constant, they find that a 10% improvement in private care over public care (i.e., less than one index point increase) would increase the number of insured people by 8.4%. In Ireland, (Harmon and Nolan, 2001) find that insurees are concerned about long waiting times. Asking some 1,100 insurees their reasons for having health insurance, 86% rate '[b]eing sure of getting into hospital quickly when you need treatment' as very important. They also find that accessibility is the most important factor motivating insurance purchase.

Other studies identified the cost of waiting through willingness to pay for quicker admissions. There is a cost in waiting as a good is worth less today if its consumption is delayed (Lindsay and Feigenbaum, 1984). Using a stated preference technique, which asked respondents their most likely action in a hypothetical scenario described by attributes, Leung et al. (2004) find that patients who value time highly choose private treatment that is readily accessible and are prepared to pay to reduce waiting time to treatment. Propper (1990; 1995) and Johannesson et al. (1998) also find evidence that individuals are willing to pay non-trivial amounts to avoid waiting for medical treatment.

There is no consensus in the literature on whether waiting lists can be used as a proxy for waiting times. Waiting lists may reflect advances in technology that permit more procedures to be done in a given time; population growth and changes in demographic composition may lead to increased demand for particular treatments (e.g., old-age related diseases); political pressure may focus on reducing the number of individuals waiting beyond a specified time without reducing overall waits. Sobolev *et al*. (2004) find a positive association between list length and waiting time, especially when the waiting list is long.[3] Sobolev *et al*. (2006) has similar findings among patients queuing for coronary artery bypass surgery: a very long list (over 60 patients) can lead to delays in treatment of a magnitude of three to four times the waiting time of the shortest list (0-14 patients). On the other hand, there are plausible reasons why the size of the waiting list may not relate to access to treatment. For example, perception of a long waiting list may deter patients and physicians from initiating surgical treatment and so deter patients from entering the waiting list. Also a long list does not translate to a long wait if patients are processed quickly.

[3] Waiting list data are defined at the patient-level. For each patient, the list length measures the number of patients with a higher or equal urgency category already on the list at registration.

Siciliani (2008) shows that the waiting list is a different concept to waiting time and these two variables can move in different directions. He uses two concepts of demand: potential demand, which is the addition to the waiting list if there is no waiting time, and current demand, which is affected by both potential demand and current waiting time. If supply grows faster than the potential demand, both waiting time and the waiting list would fall in the short run as supply is higher, but in the long run, the increasing supply will induce higher demand, which implies a higher stock of patients on the waiting list. Waiting time however may still be falling. Several studies have found that the waiting list can remain constant while waiting times improve. Using Canadian administrative data on several surgical procedures ranging from coronary artery bypass graft to knee replacement, DeCoster *et al.* (1999) find that while the waiting list increased, waiting times for these procedures remained stable, or reduced. Martin et al. (2003) find no evidence that long waits are linked to hospital capacity, which determines queues. Newton et al. (1995) use seven-year quarterly data from the English National Health System and find evidence of a dynamic effect, in which increased admissions improve waiting times but the list size remains unchanged because the list expands at the same time.

3 DATA AND METHODOLOGY

This study uses two sets of data: hospital administrative data for waiting times also used to construct waiting lists, and a national household survey containing insurance information and socio-demographic controls. The hospital data are derived from inpatient episodes and matched waiting time data from public hospitals in NSW consisting of patients waiting for planned procedures who *completed* their hospital stay between period 1/07/2004 and 30/06/2005. It records the date the patient was placed on the hospital's waiting list by the specialist (listing date)[4] and the admission date of the patient to the hospital (removal date). The patient waiting time is defined as the duration between listing and removal dates. For the individual data we use the Household, Income and Labour Dynamics in Australia (HILDA) survey for the year 2005/2006. HILDA is a panel study of Australian households. It contains information on private health insurance holding, income, education, and household characteristics. The HILDA data are linked with the hospital data through postcode of residence of observations in the HILDA sample.

[4] Excluding any outpatient waiting period.

Waiting lists are counts at a specific date. The waiting list is "backward looking" as it captures patients who entered the waiting list at time (t – k) (where k =1, 2, …) and remained untreated at time t. To construct waiting lists from our hospital data, we define a census date. We specify 1/07/2004 as the census date, so the waiting list is the count of patients, admitted during 2004-05, who are on the list at the end of June 2004. All patients in our data listed after 1/07/2004 are excluded from the constructed waiting lists. For consistency with waiting lists constructed from aggregate data, which are unlikely to contain information about patient's characteristics, we place no restrictions on age, payment status, and citizenship status of the patients on the waiting list.[5] Table 1 shows the distribution of the waiting list by listing year and admission year. The earliest listing year is 1996. The table shows that a large majority of listed patients get admitted quickly and that the proportion of patients treated during 2004-05 declines with duration on the list.[6] Almost 90% of patients were admitted within a year of listing date. This suggests that potential bias from omission of patients on the list on the census date who were admitted to hospital after the period of our data (30/06/2005) is minimal. In total, there are about 60,000 people on the waiting list, and this figure is consistent with that reported by the government's Audit Office (NSW Audit Office, 2003).

We undertake our analysis for 16 Area Health Services (AHS) and 575 postcodes within NSW.[7] The aggregation by AHS is comparable to the regional health authorities used by BHP. On average, there are about 9 patients on the waiting list per thousand of population and 1 patient on the long-term waiting list per thousand of population. The mean waiting time is 83 days. In constructing the waiting time variables by AHS, we include all patients in the hospital data because this gives the most accurate measures of waiting times.

The insurance variable from the HILDA data is binary, taking a value of 1 if an individual has private hospital cover, and 0 otherwise. The mean insurance rate is 53%, which is higher than the national average (45%). We control for the effect of

[5] The only exclusion is of patients who were admitted but whose status is classified as "not ready for care". Further examination of these observations reveals that the majority were admitted for renal dialysis.

[6] Dixon and Siciliani (2009) suggest an alternative way to infer the distribution of waiting list given the distribution of waiting time. Using their approach, we find that the proportion of patients on the long-term list (on the waiting list for more than 12 months) was close to zero (less than 0.3%). This is because the bulk of patients in NSW are treated within a year with a mean waiting time of around 80 days.

[7] The HILDA data contains no observations from postcodes in the most remote AHS, Far West. This AHS is therefore excluded from our analysis.

socio-demographic variations on PHI demand by including covariates similar to those used by BHP in their model of individually-purchased insurance.

BHP can include hospital characteristics of health region because they use repeated cross section data for 5 years. They find that hospital expenditures and staff characteristics are not significant predictors of insurance demand. In this paper, with a single year of data we control for differences in supply-side variables by including dummy variables for area remoteness. Detailed data definitions and summary statistics of all variables are in the Appendix. The HILDA sample size is 2,315 adult individuals (aged 18 or older).

Table 2 presents means of all variables in the insurance demand model and Table 3 reports insurance rates and waiting list and time variables by AHS. The insurance rate is 52.7% and varies quite widely by AHS from 29% in Macquarie to 78% in Northern Sydney. On average there are 9.4 and 1.0 person per ‘000 population on the total and long term waiting list respectively. Waiting list variables exhibit wide variation by AHS. For example, in two of the sixteen AHSs (Central Coast and Mid Western), there are 2 patients per thousand population on the long term list, whilst in all other areas the long term list is smaller. In three AHSs 10% of patients waited less than 4 months to be admitted, whilst 10% of patients in four AHSs waited at least 6 months to be admitted. The average waiting time is 82.7 days and 5.3% have waits in excess of one year. The longest mean waiting time is 127 days and in this AHS 10% wait more than 1 year. The average wait at the 90th percentile is 145 days, and this varies markedly by AHS, from 72 days in Northern Sydney to 272 days in Illawarra.

Long-term waiting lists are positively and significantly correlated with the proportion of long wait (correlation coefficient of 0.75 (p-value=0.05), whilst there is no significant pairwise positive relationship between waiting lists and any of the waiting time variables. Pairwise correlations among the waiting time variables are, of course, positive, high and significant.

4 RESULTS

Table 4 presents probit models of insurance choice. We present results from 5 independent models with varying covariates of the waiting list and time variables. Model 1 presents results that are comparable BHP’s results with total waiting lists and long-term waiting lists per ‘000 of the area’s population included as regressors.

Models 2 to 5 present results from alternative models, including different combinations of waiting list and waiting time variables.

The results in model 1 are consistent with those of BHP: insurance demand is negatively related to the total list and positively related to long-term lists. Neither of these two variables however is statistically significant. In their model of individual (not employment-based) insurance, BHP report a t-statistic of 1.84 for the long-term waiting list variable, which appears to be calculated based on standard errors of estimates that are not corrected for clustering. In our models, we report clustered standard errors which are larger than the uncorrected standard errors, as is commonly found (Moulton, 1990; Bertrand, 2004). However, even with uncorrected standard errors the long-term waiting list variable is not significant (t =1.09).

In model 2 we add waiting time variables: mean waiting time (*Mean wait*) and the proportion of patients whose waiting time exceeded 12 months (*Long wait*). Long wait is similar in spirit to the idea of long-term waiting lists, that is, it captures the upper tail of the waiting time distribution. The coefficient on *Long wait* is positive and significant at 5% level, indicating that the probability of private health insurance purchase increases with the upper tail of the waiting time distribution. This suggests that long waiting times, and not long waiting lists, matter for insurance demand.

In model 3, we test an alternative variable to capture the upper tail of the waiting time distribution: the waiting time of the upper 10% of the patient distribution (*P* 90^{th} *wait*). We find consistent results with *Long wait*, but the impact of this variable is more precisely estimated. At the mean of all variables, a 7 day increase in the *P* 90^{th} *wait* increases the probability of buying insurance by 0.03 percentage points (6% of the mean insurance rate).

In the last two models, 4 and 5, we omit both waiting lists variables but retained the waiting time variables. *Long wait* is no longer significant but *P* 90^{th} *wait* is still highly significant. An explanation for this change of result is the correlation of *Long wait* with the long-term waiting list. One needs to control for the size of the long-term waiting list to isolate the effect of *Long wait*. Using the Bayesian Information Criterion, the preferred model is model 5. In this model, a 7-day increase in the P 90^{th} *wait* increases the probability of buying insurance by 0.02 percentage points (4% of the mean insurance rate).

While measures of long waiting time suggest a positive relationship with private insurance demand, the average waiting time (*Mean wait*) tends to lower insurance demand. Computing the marginal effects of *Mean wait*, we find that they are larger in magnitude than those of the long wait variables suggesting that the overall effect of waiting time on insurance demand is quite small (due to the counteracting effects of average and long waiting time), or even negative.[8] Waiting time variables move together, however for small increases, long waiting time may increase without changing the average wait.

The coefficients of individual characteristics are stable across models. The probability of insurance purchase increases with income, education, and age. Those who are affected by the Medicare Levy Surcharge (MLS), which penalises high-income earners for not having private health insurance, are, as we expected, more likely to buy insurance. Home owners and people living in the cities are also more likely to have health insurance. On the other hand, households with small children and larger households are shown to be less likely to buy insurance. As we have controlled for age and household income, household composition may capture other determinants of the insurance decision that relate to tastes or availability of other smoothing mechanisms besides insurance.

[8] We are not suggesting that the effects of average wait and long waiting time should be added together. However, to illustrate our point, let ρ be the regression coefficient of a simple regression of long waiting time on average waiting time. In the data, $\rho = 2.4$. We can rewrite an insurance model as a function of just one the variable (say, average waiting time). The coefficient of the average waiting time will be the sum of its own effect and the effect of long waiting time, scaled by ρ. Our estimation results (model 5) find that the first term is negative and the second term is smaller positive, but would have to be scaled up by ρ. The resulting coefficient is very small due to the counteracting effects.

5 DISCUSSION

Anecdotal evidence from many countries suggests that waiting times for elective procedures in public hospitals motivate people to purchase PHI to access private treatment. However there is very little empirical evidence for this using waiting time data. In the absence of waiting times, the size of public hospital waiting lists and particularly the list of those waiting more than 12 months, has been used as a proxy for waiting times in models of insurance demand.

Probably the most influential paper is Besley et al. (1999) which finds large impacts on insurance purchase from reducing the long term list. This has resulted in recommendations to introduce policies to increase private insurance cover to reduce public hospital waiting times. Australia introduced such a policy, a 30% premium subsidy, in 1999 with the aim of reducing pressure on the public hospital system. There have been recommendations for similar policies in the UK (Siciliani, 2008).

The main objective of this paper has been to test the usefulness of waiting lists as a proxy for waiting time in relation to demand for private health insurance. Unlike BHP, we do not find that the long list is a significant determinant of demand for insurance. However we find that long waiting times do significantly increase insurance. This suggests that the relationship between waiting times and waiting lists is not as straightforward as is commonly assumed.

The large impact on individual PHI purchase of the long waiting list in the BHP results is only marginally significant, and the significance may not be robust to standard error corrections for clustering by regional health authority and year. Furthermore, a close examination of the BHP analysis suggests that their results may not reflect a systematic long-run relationship between PHI purchase and the size of a long-term waiting list. In the model with time and area fixed effects, it seems that BHP's identification of the long-term list effect relies on the presence of areas where long-term waiting lists change faster than average . As such, their model predicts that people's insurance decisions depend on relative changes in long-term waiting lists, i.e. people would drop their insurance coverage if their area's long-term waiting list decreased by more days than in other areas. This co-variation is different from the relationship between the level of the waiting list and PHI coverage.

Also, the analysis of BHP could be sensitive to data construction. In the final year of BHP's analysis (1991), there was a large fall in the long-term waiting list, which had been stable in the previous 5 years (1986-1990). In addition, the insurance rate, which had been trending upwards since the beginning of the study period, reaching its highest level in 1990, also fell in 1991. If most of the changes in the area-level waiting list and insurance coverage used to identify the effect of the waiting list on insurance occurred around this time, it is possible that BHP's results are sensitive to their choice of study period. A policy shock which lead to a shorter long-term waiting list, could drive the positive relationship found between the long-term waiting list and the insurance rate.

Our finding, that that long waiting times increase the demand for insurance, should be interpreted with caution because the relationship between waiting times and insurance demand is complex. The negative and significant impact of average waiting times tends to offset the effect of waiting times in the upper tail of the distribution. These measures are highly correlated so that policies that target the upper tail of the distribution will impact on the mean as well, with perhaps unintended impacts on insurance coverage.

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Table 1: Waiting list observations by listing year and removal year

	Admission year		
List year	2004	2005	Total
1996	1	0	1
1997	1	1	2
1998	1	3	4
1999	9	26	35
2000	44	20	64
2001	184	108	292
2002	1,167	590	1,757
2003	9,677	4,119	13,796
2004	33,590	8,030	41,620
	44,674	12,987	57,661

Table 2: Variable means

Variable	Mean	Std. Dev.
Insurance	0.527	0.499
Total List	9.235	3.822
Long-term List	0.947	0.663
Mean wait	82.683	21.525
Long wait	0.053	0.025
P 90^{th} *wait*	145.277	52.453
Income	81.654	70.016
MLS	0.326	0.469
Postgraduate	0.049	0.216
Graduate	0.065	0.247
Undergraduate	0.146	0.353
Diploma	0.321	0.467
High School	0.108	0.311
Owner-occupier	0.690	0.462
Male	0.418	0.493
Age 30s	0.179	0.384
Age 40s	0.216	0.412
Age 50-65	0.273	0.446
Age 65+	0.201	0.401
Children	0.492	0.924
Adults	2.129	1.003
City	0.496	0.500
Inner region	0.362	0.481
Outer region	0.143	0.350

Table 3: Insurance and waiting list variables by AHS

AHS	N	PHI cover	*Total List*	*Long-term List*	*Mean wait*	*Long wait*	*P 90th wait*
Central Coast	158	0.544	14.560	2.566	96.842	0.089	197
Central Sydney	154	0.591	6.862	0.601	63.848	0.036	99
Greater Murray	96	0.385	9.723	0.740	85.173	0.048	147
Hunter	247	0.449	6.697	0.416	68.791	0.032	109
Illawarra	169	0.556	15.205	1.820	127.001	0.099	272
Macquarie	45	0.289	13.705	1.259	82.775	0.044	132
Mid North Coast	106	0.406	14.716	1.319	116.342	0.082	231
Mid Western	89	0.382	14.622	2.146	78.893	0.057	125
New England	105	0.381	13.425	0.500	71.506	0.024	130
Northern Rivers	148	0.412	4.800	0.578	96.141	0.074	150
Northern Sydney	275	0.775	4.522	0.241	46.308	0.019	72
South Eastern Sydney	241	0.614	6.629	0.793	79.477	0.056	136
South Western Sydney	162	0.506	11.225	1.246	103.908	0.074	180
Southern	77	0.442	10.588	0.371	80.468	0.023	146
Wentworth	88	0.466	6.985	1.062	87.689	0.057	134
Western Sydney	155	0.594	7.597	0.680	78.681	0.044	134
Total	2135	0.527	9.235	0.947	82.601	0.053	145

Table 4: Probit coefficients with waiting list variables

	Model 1	Model 2	Model 3	Model 4	Model 5
Waiting					
Total List	-0.008	0.049	-0.046*		
	(-0.38)	(1.47)	(-1.73)		
Long-term List	0.068	-0.327	0.126		
	(0.57)	(-1.45)	(1.01)		
Mean wait		-0.019**	-0.026***	-0.007	-0.019***
		(-2.23)	(-3.54)	(-1.46)	(-2.91)
Long wait		17.83**		5.979	
		(2.15)		(1.37)	
P 90^{th} wait			0.012***		0.008***
			(3.42)		(2.86)
Individual					
Income	0.009***	0.009***	0.009***	0.009***	0.009***
	(5.46)	(5.33)	(5.39)	(5.31)	(5.54)
MLS	0.270**	0.282**	0.276**	0.277**	0.265**
	(2.11)	(2.21)	(2.19)	(2.17)	(2.10)
Postgraduate	0.510***	0.498***	0.500***	0.501***	0.516***
	(3.14)	(3.03)	(3.04)	(3.07)	(3.15)
Graduate	0.476***	0.474***	0.474***	0.474***	0.492***
	(3.21)	(3.18)	(3.18)	(3.20)	(3.30)
Undergraduate	0.715***	0.708***	0.716***	0.710***	0.725***
	(6.74)	(6.62)	(6.70)	(6.66)	(6.83)
Diploma	0.129*	0.134*	0.129*	0.131*	0.135*
	(1.80)	(1.87)	(1.80)	(1.82)	(1.89)
High School	0.088	0.096	0.092	0.091	0.099
	(0.75)	(0.82)	(0.80)	(0.77)	(0.85)
Owner-occupier	0.605***	0.621***	0.618***	0.613***	0.602***
	(7.29)	(7.45)	(7.45)	(7.32)	(7.18)
Male	-0.066	-0.0627	-0.0654	-0.0637	-0.068
	(-1.46)	(-1.39)	(-1.44)	(-1.41)	(-1.49)
Age 30s	0.254**	0.249**	0.258**	0.252**	0.263**
	(2.25)	(2.22)	(2.27)	(2.23)	(2.32)
Age 40s	0.299**	0.307***	0.297**	0.301**	0.305**
	(2.49)	(2.58)	(2.46)	(2.51)	(2.54)
Age 50-65	0.650***	0.646***	0.648***	0.651***	0.653***
	(5.38)	(5.39)	(5.42)	(5.38)	(5.41)
Age 65+	0.630***	0.624***	0.617***	0.626***	0.637***
	(4.62)	(4.60)	(4.52)	(4.57)	(4.63)
Nm children	-0.118**	-0.112**	-0.121**	-0.114**	-0.122**
	(2.30)	(2.21)	(2.35)	(2.24)	(2.36)
Nm adults	-0.139***	-0.140***	-0.136***	-0.137***	-0.140***
7	(3.26)	(3.26)	(3.28)	(3.17)	(3.37)
Inner region	-0.341***	-0.313***	-0.255**	-0.319***	-0.297***
	(-3.67)	(-3.42)	(-2.64)	(-3.55)	(-3.33)
Outer region	-0.291**	-0.289**	-0.125	-0.252**	-0.271**
	(-2.33)	(-2.29)	(-0.93)	(-2.30)	(-2.53)
Constant	-1.160***	-0.683**	-0.464*	-0.911***	-0.732***
	(-6.78)	(-2.44)	(-1.69)	(-3.57)	(-2.98)
Log L	-1227.6	1224.5	-1218.7	-1226.1	-1221.8
BIC	2610.1	2619.5	2607.8	2607.2	2598.5

Note: *t* statistics in parenthesis based on standard errors corrected for clustering by postcodes. The sample size is 2315. *, ** and *** indicates significance at 10%, 5% and 1% respectively.

Appendix: Waiting time and patient characteristics

Variable	Definition	Mean	Std.Dev
Waiting			
Total List	Count of patients on the public hospitals waiting list per '000 population on 1/7/2004	9.235	3.822
Long-term List	Count of patients listed for 12 months or more per '000 population on 1/7/2004	0.947	0.663
Mean wait	Mean waiting time of all patients	82.601	21.377
P 90th wait	90th percentile of waiting time	145.277	52.453
Long wait	% of patients waiting more than 12 months	0.053	0.025
Insurance			
Insurance	=1 if have PHI	0.527	0.499
Demographic			
Male	=1 if male	0.418	0.493
Income	Household annual gross income ($'000)	81.654	70.016
Nm children	Number of children	0.492	0.924
MLS	=1 if affected by the Medicare Levy Surcharge	0.326	0.469
Nm adults	Number of persons aged 15+	2.129	1.003
Owner-occupier	=1 if owned house/currently paying off mortgage	0.690	0.462
Education			
Postgrad	=1 if education is postgraduate level	0.049	0.216
Grad	=1 if education is graduate diploma	0.065	0.247
Undergrad	=1 if education is undergraduate	0.146	0.353
Dip	=1 if education is diploma/certificate III-IV	0.321	0.467
HS	=1 if education is high school	0.108	0.311
Incomplete	=1 if education is less than high school (base)	0.310	0.462
Age			
Age <30	=1 if age 18-29 years (base)	0.130	0.336
Age 30s	=1 if age 30-39 years	0.179	0.384
Age 40s	=1 if age 40-49 years	0.216	0.412
Age 50-65	=1 if age 50-65 years	0.273	0.446
Age 65+	=1 if age >65 years	0.201	0.401
Area			
City	=1 if live in city (base)	0.496	0.500
Inner region	=1 if live in inner NSW postcodes	0.362	0.481
Outer region	=1 if live in outer NSW/ remote NSW postcodes	0.143	0.350