# Eliciting the Demand for Long Term Care Coverage: A Discrete Choice Modelling Analysis

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**Summary**

We evaluate the demand for long term care (LTC) insurance prospects in a stated preference context, by means of the results of a choice experiment carried out on a representative sample of the Emilia-Romagna population. Choice modelling techniques have not been used yet for studying the demand for LTC services. In this paper these methods are first of all used in order to assess the relative importance of the characteristics which define some hypothetical insurance programmes and to elicit the willingness to pay for some LTC coverage prospects. Moreover, thanks to the application of a nested logit specification with 'partial degeneracy', we are able to model the determinants of the preference for status quo situations where no systematic cover for LTC exists. On the basis of this empirical model, we test for the effects of a series of socio-demographic variables as well as personal and household health state indicators.

**Keywords:** Health Insurance, Long Term Care, Choice Experiments, Nested Logit Models

**JEL Classification:** I11, I18, H40, C25

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# Eliciting the demand for long term care coverage: a discrete choice modelling analysis*

Rinaldo Brau[†] and Matteo Lippi Bruni[‡]

**Abstract**

We evaluate the demand for long term care (LTC) insurance prospects in a stated preference context, by means of the results of a choice experiment carried out on a representative sample of the Emilia-Romagna population. Choice modelling techniques have not been used yet for studying the demand for LTC services. In this paper these methods are first of all used in order to assess the relative importance of the characteristics which define some hypothetical insurance programmes and to elicit the willingness to pay for some LTC coverage prospects. Moreover, thanks to the application of a nested logit specification with 'partial degeneracy', we are able to model the determinants of the preference for status quo situations where no systematic cover for LTC exists. On the basis of this empirical model, we test for the effects of a series of socio-demographic variables as well as personal and household health state indicators.

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## 1 Introduction

Over the last 15 years, both the social policy debate and the economic literature have paid growing attention to the problem of ensuring adequate financing and provision of long term care (LTC) services (e.g. Eisen and Sloan, 1997). This was reflected in important reforms of the system of public benefits involving countries such as Germany in 1994, Luxembourg in 1998 and Japan in 2000. Even in countries that did not experience analogous radical

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changes, major concerns have been expressed over the increasing trends in LTC expenditures that challenge the financial sustainability of the different systems and raise delicate equity issues.

Economic literature on LTC has focused on two main issues. On the one hand, researchers have tried to provide indications on expected trends of LTC costs in order to advise policy makers on the amount of resources that are likely to be absorbed by these services. On the other hand, a great deal of theoretical and empirical work has tried to understand why the insurance market does not work effectively in transferring LTC risk. Until now, results of both strands of literature have been only partially satisfactory.

As for future LTC costs, it is extremely difficult to develop reliable forecasts for them. Demographic and economic factors interact in a complex way, and their evolution is predictable only to a limited extent. Even very accurate projections are extremely sensitive to slight changes in the basic assumptions on the evolution of the main determinants of LTC expenditures, such as the disability rates per age group, the distribution of demand between home and residential care, and changes in the unit costs of care (Hancock *et al*, 2003). At the same time, it has been shown that this phenomenon is often associated with a reduction in the difference between female and male lifespans. As a consequence, the expected raise in LTC needs may be offset by increased opportunities of mutual support between partners (Lakdawalla and Philipson, 2002).

In principle, lack of accurate information on the demand side would not raise serious concerns, if one could simply rely on market mechanisms for ensuring an adequate coverage against the risk of disability in old age. Yet, since it is common knowledge that private insurance markets for LTC are usually very small, a series of arguments have been explored in current literature in order to find a rational explanation for this phenomenon, which is present in very different countries, irrespective of the way health and social care are organised (Norton, 2000). Potential explanations have been suggested both for the supply and demand side. In the first case, they range from market failures such as adverse selection and ex-post moral hazard which affect LTC more manifestly than the health insurance market, to the existence of aggregate undiversifiable risk (Cutler, 1993). Lack of demand is, instead, mainly attributed to the myopic behaviour of young generations, who underestimate the risk of disability; and to the "intra-family" strategic behaviour of elderly people, who foster personal care by adult children by not providing for themselves with coverage. (Pauly, 1990; Zweifel and Strüwe, 1998). Yet, a robust empirical validation for most of these conjectures remains to be found (e.g., Sloan and Norton, 1997; Sloan, Picone and Hoerger, 1997; Mellor, 2001). Besides, formal LTC services have partial substitutes that may reduce the demand for coverage. Public assistance may crowd out private demand for care or even induce individuals to strategically choose sub-optimal levels of coverage because they rely on last resort public intervention. Finally, individuals may prefer to purchase formal care at the point of demand rather than acquiring ex-ante coverage.

It has been recently shown that in the US market, although there is some evidence of supply side market failures, they do not seem to be primarily responsible for small market size (Brown and Finkelstein, 2004). The latter result confirms the importance of a better understanding of demand patterns. With this objective in mind, our paper proposes a complementary approach with respect to those usually adopted in current literature. We focus on the demand side for LTC coverage in Italy, an institutional context where it would be impossible to evaluate it from actual expenditure data, given that the market is very limited. Moreover, the problems previously discussed suggest that an unmet demand for risk coverage is likely to arise. All this paves the way to adopting a 'stated preferences' approach instead of the more common 'revealed preferences' studies. For our purpose, the advantage of studying demand in hypothetical markets is twofold. It gets rid of supply side imperfections which may hamper voluntary transactions between economic agents, and it also enables us to explore how preferences vary according to alternative institutional solutions.

Our study uses discrete choice modelling technique, and the purpose is to detect the main determinants of the demand for LTC coverage, and provide estimates for the willingness to pay (WTP) for alternative cover programmes. In particular, a choice experiment was carried out on a sample representative of the population of the Italian region Emilia-Romagna. While the use of choice modelling, since the seminal works by Propper (1990, 1995), has become standard practice in health economics, to our knowledge these techniques have never been used for studying the demand for LTC insurance. The choice modelling approach allows us to identify the relative importance of the characteristics of an insurance programme.

The answers to the choice experiment have been studied through well established regression techniques. In particular, theoretical reasons and hypothesis testing has led us to rely on a 'nested logit with partial degeneracy' specification. By doing so we have, on the one hand, endorsed and tried to make more operational the Ryan and Skatun (2004) recommendation to model the opting out option in stated preference studies. On the other hand, we have analysed both the relative attractiveness for different insurance prospects, and the overall propensity to insure against LTC-related expenditures, i.e. the choice of whether to prefer one of the hypothetical policies proposed, or the existing situation. In view of this, we test for the effects of a series of socio-demographic variables, family composition, personal and household health status indicators. We find a strong significance for the selected attributes in determining the WTP, with indications in line with economic theory. Also demographic and personal status indicators display clear significance in modelling the opt-out (i.e. choice of the status quo) stage in the nested logit framework.

The structure of the paper is the following. The next section illustrates the main features of the survey carried out, while section 3 is devoted to the presentation of the econometric framework. In section 4 we present the results of our estimates. Finally, section 5 concludes and discusses the policy implications of WTP estimates.

## 2 The dataset and the discrete choice experiment

Italy is characterised by the absence of universal programs for covering LTC expenditures. In order to collect information on the potential interest for implementing new programs aimed at financing LTC expenditures, a survey was carried out between October and December 2002 on a regional scale as part of a national interest project on LTC. A questionnaire collecting information on socio-economic status, health conditions and household demographic composition was submitted by means of 1405 personal interviews to a representative sample of the population of the Italian region Emilia-Romagna. In this paper we use a subsample of 1176 observations for which the information about the net family income is available, and for which a check was run on the internal consistency of the answers.

The interview included a discrete choice experiment designed to elicit the WTP for LTC coverage. Originally developed in transportation and marketing literature, the choice experiment technique has increasingly found applications in environmental economics (e.g., Hanley, Mourato and Wright, 2001) and, in more recent years, in health economics (e.g. Propper, 1995; Ryan and Gerard, 2003).[1]

In the setting-up of a choice modelling analysis, the first issue to be addressed is the definition of a hypothetical scenario that serves as a framework for individual choices. The scenario for this work was constructed following the indications emerged from a panel of economists and statisticians from the University of Bologna and experts of health and social services from the Regional Agency of Health Care Services of Emilia Romagna. The choice of the attributes was based on two main criteria: policy relevance and feasibility of administration addressed within general personal interviews carried out in respondents' homes. In order to reduce task complexity for respondents, no more than four attributes were considered.

Indeed, the definition of the scenario is typically a very critical operation, and we have to consider that LTV encompasses a wide range of services and includes levels of disability that vary considerably. Moreover, because for the same health conditions different transfer schemes can be designed, ranging from in-kind provision of care, to cash payment defined according to severity or to the expenses actually afforded. The survey tackles this complexity by anchoring the insurance coverage to a specific health status, described as a condition in which people need help for several hours per day in their activities of daily living and for which both home and residential care can be considered appropriate from a clinical point of view although they are different with respect to the monetary cost and the burden of care-giving left to the family. In order to ensure a homogeneous perception of the health status described above, the amount of care needed was also quantified in monetary terms, by prospecting a monthly cost of 1550 euros (former 3,000,000 ITL) in the case of residential

[1] While we are not aware of oher choice modelling studies on the demand for LTC insurance, a few analyses based on contingent valuation methods have recently appeared (Costa-Font and Rovira-Forns, 2004; Brau *et al*, 2004))

care and of 1033 euros (former 2,000,000 ITL) for home-care. It was specified that these amounts had to be considered as extra-costs, in addition to the support currently offered by the public sector. Moreover, respondents were informed that the service proposed did not imply the lack of coverage for more (or less) severe conditions.

A second problem is represented by the typical form assumed by existing health insurance schemes, which usually include clauses for the extension of coverage to family members. The extension to additional members of the household is straightforward in case of public coverage, where the service typically covers the entire population. In case of private voluntary insurance, extensions are usually available in standard contracts. However, had the survey allowed for that, the individual WTPs recorded would have been referred to inherently different goods since each respondent could have thought of (implicitly) covering a different number of people according to each person's family conditions. So as to circumvent this problem, respondents were explicitly informed that the insurance plans proposed in the choice experiment were to be considered as covering only the respondent, notwithstanding the existence of a wider range of possibilities in the real world.

Starting from this common framework, some hypothetical insurance schemes for LTC risk were proposed to the respondent. Each alternative varied with respect to the values and characteristics assumed by four relevant attributes: a) the funding scheme, b) the insurance premium, c) the degree of coverage (copayment rate), d) the option right for extending coverage to the additional expenditures determined in the case residential care is chosen. The following table describes the attributes and related levels which were used in the choice experiments.

Table 1: *The attributes and levels for the LTC policy*

| **Attributes** | **Levels** |
|---|---|
| Financing scheme: | *Public* (general taxation/compulsory participation) |
| | *Private* (insurance premium/voluntary participation) |
| Yearly cost of coverage: *(in Euros)* | *103 258 387 516 775* |
| Degree of coverage : | *Low coverage (70% copayment rate)* |
| | *Medium coverage (50% copayment rate)* |
| | *High coverage (25% copayment rate)* |
| | *Total coverage (0% copayment rate)* |
| Option for covering additional costs of residential care: | *Included* / *Not included* |

For the case of public funding, it was expressly stated during the interview that the proposed solution consisted of a homogeneous coverage provided to the whole population. Participation was to be considered compulsory, and the service financed by means of an increase in the income tax actually paid by the respondent. Respondents were informed that the price indicated as the insurance cost represented their own additional tax. This typically implies that richer than average respondents would be asked to contribute more, and vice-versa for the poorer. Conversely, in case of private insurance participation was voluntary, and the level of coverage could vary according to the subscriber's preferences. Moreover, each subscriber would pay a premium that depends on the features of the insurance policy and to some extent on his personal conditions.

The option for covering additional costs consisted of applying the co-payment rate to the entire amount of expenditures, also in cases where the subscriber opted for residential LTC provision. When this later option was not included, the policyholder could still opt for residential care, but had to bear the entire amount of the consequent additional costs. In particular, he would not receive any reimbursement for the extra expenditure incurred from choosing nursing home care.

The combination of these attributes and their levels yields a full factorial of 64 possible alternative insurance packages. We selected half of this factorial using an end-point fractional design, so as to allow for interactions between the extremes of the attributes (see Louvière *et al*, 2000). We also introduced a "status quo option", consisting of no additional coverage relative to the level guaranteed by the public sector at the time of the interview. As is well known, in some cases the "no choice" alternative has no explicit economic meaning. In this case though, the choice of the status quo implies that respondents prefer the current level of support and are not willing to extend coverage. Figure 1 below presents an example of one of the show-cards employed to submit various choice sets during the interviews. Each respondent was asked to select his preferred alternative for 11 different choice sets.

Figure 1: An *example of the show-cards used in the survey.*

Let us assume that only the three solutions below are made available. Which one would you choose?

| **Characteristics of the service** | **Solution A** | **Solution B** | |
|---|---|---|---|
| ***Way of payment*** | Public coverage | Private insurance | **Present situation** |
| ***Copayment rate*** | Total coverage (0% left to the patient) | High coverage (25% left to the patient) | |
| ***Option for covering additional costs*** | Without the option to cover residential costs | With the option to cover residential costs | |
| ***Cost of the coverage*** | Lire 500.000 (€ 258) per year | Lire 1.500.000 (€ 775) per year | |
| **Preference** (thick only one) | ❒ | ❒ | ❒ |

In order to control for the respondent's actual understanding of the exercise, one of the choice sets contained a (strictly) dominant alternative, i.e. a profile which has the same qualitative attributes (public or private financing and the possibility or not to extend coverage to the additional costs of residential care) and does better for at least one of the quantitative attributes (cost and copayment rate) and does not do worse for the others. The (very few) respondents who chose the dominated solution were excluded from the sample. For the respondents who made the "correct" choice, the dominant card was excluded from the estimation, since the decision on that choice set could not be considered informative with regard to the trade-off between attributes.

In experiments with several choices, it is unlikely that replications from the same individual will be truly independent, and order bias may arise. To limit this possibility, the show-cards with the experimental choice sets were rotated sequentially. In general, the only effects with repeated choices are on statistical efficiency, but not on unbiasedness (Louvière *et al* (2000, ch. 9). Train (2003, p. 46, 55) observes that logit probability estimations can handle the dynamics of repeated choice, including state-dependence, inasmuch as unobserved factors that affect decision are independent over time in the repeated choice.

The sample used for the estimation also includes individuals (23% of the sample) who have always chosen the status quo. We checked whether this could significantly affect the quality of the results. More precisely, with an ad hoc question, we identified those respondents who actually did not consider at all the possibility to choose one of the insurance alternatives proposed since they were not interested in any LTC cover scheme at all. We found that these answers were quite uniformly distributed throughout the sample, without significantly changing the distribution of the design and other relevant characteristics, such as family income, respondent's age and education level.

Starting from respondents' stated choices, the choice modelling approach enables us to evaluate the service on a monetary metric basis, under the assumption that the overall utility equals the sum of the utilities obtained from each attribute. We conjecture that utility decreases with cost and copayment rate, whereas it increases with coverage extension. Conversely, there are no prior expectations on the effect of moving from a public to a private financing scheme.

## 3 The econometric approach

In most economic applications, data obtained from choice experiments have been studied by means of a model labelled multinomial logit (henceforth, MNL) by some authors (e.g. Mc Fadden, 1984), and conditional logit by others (e.g. Greene, 2003). Despite its wiespread use, the probabilistic structure of the MNL model has some implications which may prove problematic in our case; more complex approaches may be more appropriate.

In the MNL model, data arising from the $k = 1, 2, \ldots K$ mutually exclusive observed choices, and taken from a sample of $h = 1, 2, \ldots, H$ respondents, can be described according to a random utility specification such as the following:

$$U\,(\text{choice } k \text{ by respondent } h) \equiv U_k^h = V_k^h + \epsilon_k^h = \boldsymbol{\beta}'\mathbf{x}_k + \epsilon_k^h, \tag{1}$$

where the elements of the vector $\mathbf{x}_k$ may either refer to the characteristics of the choice alternatives, or be individual-specific. Individual utility is given by the sum between an observable component $V_k^h$ and a stochastic unobservable component $\epsilon_k^h$. Depending on the assumptions made on the distribution of $\epsilon_k^h$, different discrete choice models are obtained. The MNL assumes that the individual random components $\boldsymbol{\epsilon}_k^h$ are independently and identically distributed (IID), with an extreme value type 1 (Gumbel) distribution with mean $\eta + \gamma/\mu$ and variance $\sigma^2 = \pi^2/6\mu^2$.[2] From the IID hypothesis, $\mathrm{cov}\left(\epsilon_k^h, \epsilon_l^h\right) = 0$, so that the variance –covariance matrix of the MNL simply reduces to $\Sigma = \sigma^2 I$.

The IID assumption leads to the so-called independence of irrelevant alternatives (IIA) property, which states that the odds of an alternative $k$ being chosen over alternative $l$ is independent of the availability of attributes or alternatives other than $k$ and $l$ (McFadden, 1984). However, when there are subsets of similar alternatives, the independence condition may prove very strong because, within these subsets, some common unobserved factors are likely to affect the error standard deviation in a common way that is different from the effect on less similar alternatives (in practice, originating different scale parameters $\mu_k$). These considerations suggest that the MNL can be not appropriated to our case, where two alternatives implying different forms of coverage extension are compared with a third solution characterised by no additional cover.

Specifying the nature of the decision process implied by our choice experiment helps understanding the problem. For any respondent, each repetition of the choice experiment can be interpreted as the outcome of two (simultaneous) decisions:

- whether or not to extend coverage against the risk of LTC expenses or to opt for the present level of coverage;
- choice of the preferred insurance scheme between two alternatives that differ in the levels of four relevant attributes.

To be consistent with this framework, let us model the unobservables in (1) according to the following additive error structure:

$$\epsilon^h (i, j) = u_i^h + u_{j|i}^h, \tag{2}$$

where the index $j$ relates to the existing 'elementary' alternatives (*insurance A, insurance B, status quo)* and $i$ relates to the choice of whether or not to extend coverage against LTC risk. In other words, the random term affecting final choices is the sum of two independent components: a specific one (conditional on the two decisions) and a common one.

The previous additive specification is the basis for the nested logit (NL) model, where the variance (more precisely the scale parameter $\mu_i$) is allowed to differ across "nests of choices". The unobservable terms related to final choices have a Gumbel distribution with variance

[2] The symbol $\eta$ indicates the mode of the distribution, $\mu$ is a positive scale parameter and $\gamma$ is Euler's constant equal to 0.577.

$$Var\left(u^h_{j|i}\right) = \frac{\pi^2}{6\mu_i^2},\ \forall i, j, \tag{3}$$

whilst the property of an equal variance is kept within each nest (in our case determined by the decision of whether or not to choose a cover against LTC risk):

$$Var\left(u^h_i\right) = \frac{\pi^2}{6\lambda^2}. \tag{4}$$

The NL model represents the most usual technique used when standard testing procedures reject the IIA assumption. By partitioning the overall decision process according to the two (or more) choices, NL keeps the IID condition of the error terms within each partition, whereas the $\epsilon^h(i,j)$ are correlated within nests. As outlined for example by Hunt (2000), organising alternatives in clusters reflects a supposed similarity between "grouped alternatives". Individuals are hypothesised to consider these as more similar than the alternatives placed from different clusters. It is because of this structure that the use of NL models has been advocated for the analysis of those choice decision cases where the possibility of "non participation" exists (e.g. Morey, 1999). In health economics literature, a similar approach has been proposed by Ryan and Skatun (2004), for analysing the 'opting out alternatives' in discrete choice experiments. Similarly, in our study the status quo alternative is actually a non participation alternative, which is intuitively different from making a choice among the insurance alternatives.

By framing the abovementioned two choices as the two nests of a two-level NL model, we end-up in a 'NL with partial degeneracy', given that there is only one single 'no insurance' option. In the case of our choice experiments, in the first nest the respondent chooses whether or not to extend currrent LTC coverage; in the second one he selects his preferred alternative conditional on having chosen to insure or not.

## 3.1 *A nested logit model with partial degeneracy*

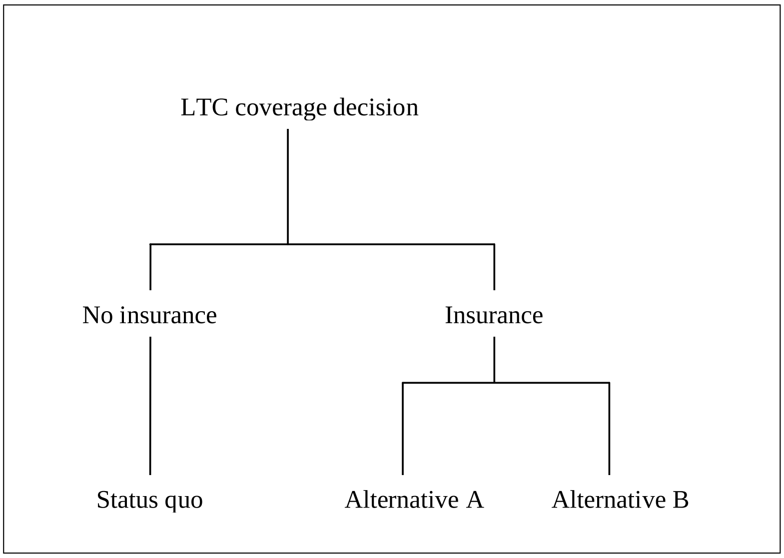
It is well known that socio-economic variables need to be considered as alternative specific, unless the usual linear-in-utility structure is abandoned (e.g. Cherchi and Ortúzar, 2003). Deciding to include individual specific variabless only in one alternative often is somewhat arbitrary. This is not the case, however, when, within a NL specification with a degenerate branch, the latter represents a 'non participation' alternative which reposes on a likely distinct economic rationale. In this case, a NL with partial degeneracy provides a natural framework for analysing the impact of some important individual-specific effects (such as demographic, health and economic status variables) in the utility expression for the top level (e.g. Greene, 2003; Louvière *et al*, 2000).[3]

Figure 2 represents our hypothesised decision tree. In the first stage of the decision

[3] For a rigourous analysis of nested logit models with partial degeneracy see Hunt (2000), Hensher and Greene (2002)

process the respondent chooses whether or not to buy an LTC insurance. In the second stage he makes his choice among the elemental alternatives of the choice set.

Figure 2: *The decision tree for the LTC cover choice experiment*




Here we better define the random utility structure.[4] For a generic elemental choice $j$, belonging to upper level $i$, a respondent's utility takes the form:

$$U(i,j) = U_i + U_{j|i}, \tag{5}$$

which can also be written as follows:

$$U(i,j) = V(i,j) + \epsilon(i,j), \tag{6}$$

where the $V(i,j)$ indicate the non stochastic utility components and $\epsilon(i,j) = u_i + u_{j|i}$ is the stochastic utility component.

In order to formally define our case, we distinguish between variables which influence the choice at the $j$ level ($\mathbf{x}$), and variables which affect the choice to insure or not to insure ($\mathbf{z}$). Given the hypothesis that the deterministic part of the indirect utility function is additively separable we can write

$$\begin{aligned} V(i,j) &= V_i + V_{j|i} \\ &= \boldsymbol{\gamma}'\mathbf{z}_k + \boldsymbol{\beta}'\mathbf{x}_l \end{aligned}. \tag{7}$$

By using (2), this leads to:

$$U(i,j) = \boldsymbol{\gamma}'\mathbf{z}_i + \boldsymbol{\beta}'\mathbf{x}_j + u_i + u_{j|i}. \tag{8}$$

At the upper stage (insurance decision), we define the non stochastic utility component as:

[4]Henceforth, we omit the individual index $h$.

$$U_i = \boldsymbol{\gamma}'\mathbf{z}_i + u_i. \tag{9}$$

The joint probability that household $h$ chooses alternative $j$ is given by the product between a marginal and a conditional probability:

$$\Pr[y_h = i, j] = \Pr[w_h = i] \times \Pr[s_h = j|i]. \tag{10}$$

A useful way to make the previous expression for NL models explicit is to define the probability choice system (PCS), which includes the marginal choice probabilities associated with the choice at the upper level, the conditional probabilities associated with the choices at the lower level, and the so called inclusive value (or expected maximum utility).[5] As we have seen with (3) and (4), in a two-level NL model, these variances are related to scale parameters $\lambda_i$ associated to the upper level, and to $\mu_i$ parameters for the elemental alternatives level. At the "lower" level, the conditional choice probability for the elemental alternatives can be written in the following way:

$$\Pr[s_h = j|i] = \frac{\exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)}{\sum_{j=1}^{J_i} \exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)} = \frac{\exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)}{\exp\left(IV_i\right)}, \quad \text{for all } i \tag{11}$$

where $J_i$ is the number of possible elemental choices in branch $i$ (*1 if* $i$=*not insure, 2 if* $i$ = *insure*)

The marginal probability at the "upper" level is:

$$\Pr[w_h = i] = \frac{\exp\left[\lambda_i \boldsymbol{\gamma}'\mathbf{z}_i + \frac{\lambda_i}{\mu_i} IV_i\right]}{\sum_{i=1}^{I} \exp\left[\lambda_i \boldsymbol{\gamma}'\mathbf{z}_i + \frac{\lambda_i}{\mu_i} IV_i\right]}, \tag{12}$$

where the symbol $IV_i$ defines the inclusive value:

$$IV_i = \ln \sum_{j=1}^{J_i} \exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right). \tag{13}$$

Hence the joint probability (10) takes the form:

$$\Pr[y_h = i, j] = \frac{\exp\left[\lambda \boldsymbol{\gamma}'\mathbf{z}_i + \frac{\lambda_i}{\mu_i} IV_i\right]}{\sum_{i=1}^{I} \exp\left[\lambda \boldsymbol{\gamma}'\mathbf{z}_i + \frac{\lambda}{\mu_i} IV_i\right]} \cdot \frac{\exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)}{\sum_{j=1}^{J_i} \exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)} \tag{14}$$

Two considerations are usually reported in the literature about the role of the scale parameters and the ratio $\lambda_i/\mu_i$, known as the "inclusive value coefficients" (or parameters).

[5] The formal expressions for the PCS of the NL proposed by a recent stream of literature (Hensher and Greene, 2002; Hunt, 2000) pays special attention to the peculiarities of NL with degenerate branches and the role of normalisation of the scale parameters which are associated with the variances in the nests of the model.

The first relates to the value which the IV coefficient should assume. Given: (i) that in the NL specification the variance at the lower level must be the largest since it shares (by having stochastic component $u_i + u_{j|i}$) part of its unobservables with the higher level (which has stochastic component $u_i$); (ii) the proportionality between the scale parameters of the Gumbel distribution and the standard deviation of unobservable terms, then if the NL specification is correct, the estimated IV coefficient $\lambda_i/\mu_i$ must lie in the interval $[0,1]$. This result is related to the higher degree of similarity between alternatives which share the same upper level. In fact, it can be shown that the correlation of the indirect utilities of any pair of elemental alternatives within the same nest is $\rho_i = 1 - (\lambda_i/\mu_i)^2$, which is clearly zero for $\lambda_i/\mu_i = 1$ (e.g. Ben Akiva and Lerman, 1985; Hunt, 2000). Hence, the closer the coefficient is to one (zero), the less (more) the degree of perceived similarity between the alternatives.

The second consideration refers to the identification problems entailed by the scale parameters. As we can see from equation (12), the IV parameter is identified (an estimate of the ratio can be obtained). However, this is not the case for the utility index $\boldsymbol{\gamma}'\mathbf{z}_i$, since its value is multiplied by the (unidentified) scale parameter $\lambda_i$. A similar consideration applies for the lower level utility index $\boldsymbol{\beta}'\mathbf{x}_{j|i}$, given the presence of $\mu_i$.

A normalisation of the general representation of the PCS given by equations (11-13) is therefore needed, by setting one scale parameter equal to 1 (and common to all nests). As outlined by Louvière *et al.*, there are no clear indications of the particular implications of normalizing with respect to the branch level scale parameter ($\lambda = 1$) rather than to the lower level scale parameter ($\mu = 1$). The same authors report that most empirical studies normalise the branch level utility index by setting $\lambda = 1$. From a practical point of view, this kind of normalisation has the advantage of enabling the researcher to carry out a direct confrontation of NL estimates with the parameters obtained using a MNL model; it relates normalisation to total variance of the error distribution; and leads to a simpler PCS. (Carrasco and Ortùzar, 2002). In the next section, we follow this convention, also in the light of some invariance results for the case of degenerate branches (see Hunt, 2000). The expressions of the PCS for the case $\lambda = 1$ are:

a) conditional choice probability for the elemental alternatives:

$$\Pr\left[s_h = j|i\right] = \frac{\exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)}{\sum_{j=1}^{J_i} \exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)} = \frac{\exp\left(\mu_i \boldsymbol{\beta}'\mathbf{x}_{j|i}\right)}{\exp\left(IV_i\right)}, \quad \text{for all } i$$

b) marginal probability at the "branch" level:

$$\Pr\left[w_h = i\right] = \frac{\exp\left[\boldsymbol{\gamma}'\mathbf{z}_i + \frac{1}{\mu_i} IV_i\right]}{\sum_{i'=1}^{I} \exp\left[\boldsymbol{\gamma}'\mathbf{z}_i + \frac{1}{\mu_i} IV_i\right]}, \quad \text{for all } i \tag{15}$$

c) "inclusive values":

$$IV_i = \ln \sum\nolimits_{j=1}^{J_i} \exp\left(\mu_i \boldsymbol{\beta}' \mathbf{x}_{k|i}\right), \qquad \text{for all } i. \tag{16}$$

As can be seen, the change is in the marginal probability at the branch, where the utility index is directly computable, and the $IV$ parameters reduce to $\frac{1}{\mu_i}$.[6] By using the estimate of the latter, it follows that the lower level utility index can be also identified. Notice that, given the theoretical condition $\lambda/\mu_i<1$, in this case the estimated lower scale parameter $\mu_i$ is expected to be larger than one.

## 4 Main empirical results

In this section we present the results of our estimates. We start with the outcome of a MNL estimation carried out with only the attributes as regressors. The results are reported on the left-hand-side of Table 2.

The coefficients for all attributes are highly statistically significant. From a policy perspective, the most interesting result is that public coverage emerges as the preferred institutional solution. As expected, both a decrease in the copayment rate and the option for covering the extra costs for residential care are positively evaluated. It is worth noting the relatively high value attached to the extension of the coverage to residential care expenditures, which is perceived as a serious risk. By computing the ratios between the estimated coefficients we can get an indication of the relative importance of the various attributes. The option for residential care is evaluated as much as 31.3 percent point of coverage, and obtaining public coverage instead of a private one is valued 16.6 percent coverage points. However, the McFadden-Hausmann test indicates (chi-sq. = 650) a strong violation of the IIA hypothesis. Although based on a regression with the design attributes only, this test is fully reliable, given that the variation of individual-specific attributes, which do not vary among the choices, does not affect the stated choices.

A way of overcoming the restrictions imposed by the IIA hypothesis is by means of a simple "random effects" multinomial probit estimation (though more structured autoregressive specifications could also be considered), which results are reported on the right-hand-side of Table 2. Once accounting for the implicit different scale parameter, the explanatory power of the attributes is confirmed and no large variations emerge for the relative size of the parameters (for instance the option for residential care and the differential effect of public coverage are now evaluated as much as 28.5 and 16.3 percentage coverage points, respectively). A likelihood-ratio test carried out according to suggestions made by Greene (2002) strongly confirms the violation of the IIA hypothesis.

*Table 2: Multinomial logit and multinomial probit estimates*

[6] This is true for the non degenerate partition, whereas for the degenerate branch it is not identified: see Hunt, 2000.

| *Variable* | *"Main effects" MNL model* | | | *"Main effects" MNP model* | | |
|---|---|---|---|---|---|---|
| | *Coefficient* | *t-stat.* | *Prob* | *Coefficient* | *t-stat.* | *Prob* |
| Financing scheme (0 private, 1 public) | 0.2960 | 11.26 | 0.000 | 0.1578 | 9.70 | 0.000 |
| Extension to residential care expenses | 0.5564 | 21.31 | 0.000 | 0.2766 | 12.85 | 0.000 |
| Degree of percentage coverage | 0.0178 | 31.20 | 0.000 | 0.0097 | 15.24 | 0.000 |
| Yearly cost of coverage | -0.0016 | -25.91 | 0.000 | -0.0008 | -14.05 | 0.000 |
| Alternative specific constant (0=status quo) | -1.3933 | -25.81 | 0.000 | -0.5683 | -8.09 | 0.000 |
| *Diagnostic statistics and tests* | | | | | | |
| Log likelihood function | -11625.3 | | | -11585.7 | | |
| Pseudo R-squared | 0.100 | | | 0.103 | | |
| IIA Tests | Mc Fadden/ Hausman test (Excluded choice "status quo"). | | | Likelihood ratio test for IIA | | |
| | Chi-Squared 650.2 | | | Chi-Squared 70.2 | | |
| Number of observations | 11760 | | | 11760 | | |
| *(Number of respondents)* | *(1176)* | | | *(1176)* | | |

We move consequently to the estimates of the NL specification described in the previous section. With a partially degenerated NL, all attributes of the degenerate alternative as well as all the observation-specific effects in a utility expression for the top level of the decision tree. In our case, there are no specific attributes related to the status quo alternative, so that we simply relate the whole vector $z_l$ of individual characteristics to the choice between extending coverage or maintaining unchanged the level of protection ensured by current public support for disabled elderly people.

NL estimations reported in Table 3 have been carried out by normalising on $\lambda_i = 1$, for all $i$, and this permits a direct comparison with the results of the MNL specification. The right-hand-side of the table reports results by a model where the endpoint design of the experiment has been exploited. Namely, to the regressors of the main effects design we added the interactions between public coverage and cover extension for the extreme cases of a 0 % and 70 % copayment rate. Only the statistically significant interaction terms are considered.

A wide range of socio-demographic indicators have been successfully used to define the index $\boldsymbol{\gamma}'\mathbf{z}$. Following most examples in the literature, they have been included as determinants of the 'no insurance' choice, so that the associated coefficient represents the (negative of the) differential effect of the socio-demographic variables on the utility of extending coverage as compared to maintaining the status quo.A first important remark is that the value of the IV parameter (0.549 and 0.566) and its significance indicate the statistical appropriateness of adopting the two level NL specification, since the IV coefficient should lie between 0 and 1, where the latter value indicates that a MNL specification should be preferred. Overall, it can be seen that a large share of the regressors display significant effects and the sign of the coefficients meets prior expectations in most cases. At the top of the table, the four attributes included in the insurance package are all highly statistically significant, although lower t-statistics indicate higher standard errors than in the MNL model. Note also that the coefficients have smaller size in absolute value because of the higher common scale parameter implied by the NL. The coefficients of the continuous variables (cost and copayment rate) both have the expected sign. The same holds for the option to extend cover to additional residential care expenses (dummy equals 1 when the option is included). We leave a more detailed analysis of the relative importance of various attributes to the next subsection, where monetary equivalents of the estimates are presented.

In reviewing the results related to the decision of whether or not to insure, recall that a positive coefficient in our model indicates that there is a higher probability that the status quo will be opted for. Our specification first includes a group of socio-demographic variables such as household income (defined as net monthly household income), respondent's age, household size, presence of adult and young children in the household.

Income positively influences the probability of extending coverage (i.e. negatively affects the choice of the status quo). This result is to some extent in contrast with previous evidence obtained from revealed preferences studies on the US private LTC market (Sloan and Norton, 1997) and suggests that at low income levels individuals may have priorities other than LTC coverage. Given the peculiar nature of LTC, policymakers may negatively evaluate the possibility that citizens would acquire substantially different degrees of cover due to differences in income, even if this outcome reflect household preferences. If the policymaker's objective function incorporates some specific egalitarian argument for elderly care, the result for the income coefficient indicates that contributions to public programs should be designed in a rather progressive manner, possibly including exemptions for low income groups. This would be the only way to provide coverage to low income groups without imposing a constraint that would divert part of their budget towards a service which they not consider a priority. Contrariwise, tax allowances on private policies, which are effective in meeting individual preferences, are likely to widen the difference in the level of protection among different income groups.

The inclusion in the regression of a dummy variable for the presence of adult children living at home allows for an assessment of the "intra-family moral hazard hypothesis" (IMH). It has been argued that since elderly people prefer to be assisted by their family members, they may strategically choose not to insure (e.g. Pauly, 1990; Zweifel and Struwe 1998). Consequently, adult children have an incentive to provide informal care for their parents, given that the entire amount of money spent in formal care would reduce future bequests. In our model, the presence of adult children decreases the likelihood of choosing a larger coverage, but the effect is not significant. In analogy with empirical works for the US market (Mellor, 2001; Sloan and Norton, 1997), our estimates do not confirm the conjecture. The IMH hypothesis does not seem to have a relevant influence on household decisions over the demand for LTC coverage even when only demand side effects are considered.

The result relating to the age variable suggests that younger generations are more favourably oriented towards an extension of cover. The myopic attitude of young people

- who would not demand cover due to a severe underestimation of their future risk of disability - has often been suggested as an explanation for the lack of demand of LTC insurance.

*Table 3: Estimates results with the nested logit specification*

| *Explanatory variables* | ***Nested Logit with main effects*** | | | ***Nested Logit with design interactions*** | | |
|---|---|---|---|---|---|---|
| | *Coefficient* | *t-stat.* | *Prob* | *Coefficient* | *t-stat.* | *Prob* |
| *"Choice of alternatives" process* | | | | | | |
| Financing scheme (0 private, 1 public) | 0.1930 | 9.80 | 0.000 | 0.2443 | 9.49 | 0.000 |
| Extension to residential care | 0.3413 | 12.88 | 0.000 | 0.3900 | 12.50 | 0.000 |
| Degree of % coverage | 0.0120 | 14.78 | 0.000 | 0.0116 | 12.77 | 0.000 |
| Yearly cost of coverage | -0.0011 | -14.50 | 0.000 | -0.0011 | -14.48 | 0.000 |
| Interaction between "extension" and "low coverage" | | | | -0.2176 | -4.23 | 0.000 |
| Interaction between financing scheme and "total coverage" | | | | -0.1083 | -2.59 | 0.010 |
| ASC (0 for status quo) | -1.1563 | -4.81 | 0.000 | -1.1388 | -4.71 | 0.000 |
| *"Insurance decision" process* | | | | | | |
| Age | 0.0195 | 7.93 | 0.000 | 0.0195 | 7.93 | 0.000 |
| Family Income in € | -0.0002 | -6.43 | 0.000 | -0.0002 | -6.41 | 0.000 |
| Sex (1 if male) | -0.1597 | -3.63 | 0.000 | -0.1601 | -3.64 | 0.000 |
| Household size | 0.1903 | 7.86 | 0.000 | 0.1903 | 7.85 | 0.000 |
| Spouse | -0.0050 | -0.09 | 0.930 | -0.0041 | -0.07 | 0.943 |
| Young children | -0.2302 | -4.06 | 0.000 | -0.2308 | -4.07 | 0.000 |
| Adult children | 0.1119 | 1.33 | 0.185 | 0.1112 | 1.32 | 0.188 |
| University degree education | -1.1502 | -6.39 | 0.000 | -1.1457 | -6.36 | 0.000 |
| Secondary school education | -0.6777 | -4.00 | 0.000 | -0.6742 | -3.98 | 0.000 |
| Compulsory education | -0.5778 | -3.48 | 0.001 | -0.5749 | -3.46 | 0.001 |
| Blue collar occupation | 0.1049 | 1.42 | 0.154 | 0.1070 | 1.45 | 0.146 |
| White collar occupation | -0.0270 | -0.42 | 0.673 | -0.0264 | -0.41 | 0.680 |
| Retired | -0.1988 | -2.76 | 0.006 | -0.1978 | -2.74 | 0.006 |
| Not working | 0.0272 | 0.35 | 0.725 | 0.0290 | 0.38 | 0.707 |
| Other employment status | 0.4555 | 3.12 | 0.002 | 0.4578 | 3.13 | 0.002 |
| Chronic disease | -0.0491 | -0.90 | 0.368 | -0.0502 | -0.92 | 0.358 |
| Self assessed health status (0 for good, 1 for bad) | 0.3342 | 6.28 | 0.000 | 0.3351 | 6.29 | 0.000 |
| Subscriber of a private health insurance | -0.4719 | -8.11 | 0.000 | -0.4714 | -8.10 | 0.000 |
| In hospital in the last year | 0.0629 | 0.94 | 0.345 | 0.0631 | 0.95 | 0.344 |
| Smoker | 0.0782 | 1.66 | 0.097 | 0.0790 | 1.68 | 0.094 |
| Preference for "cash" LTC coverage | -0.0203 | -0.49 | 0.624 | -0.0215 | -0.52 | 0.602 |
| Existence of a person with LTC disability in the family | -0.2535 | -5.17 | 0.000 | -0.2531 | -5.16 | 0.000 |
| Health in the first 3 priorities for new public expenditures | -0.3229 | -6.25 | 0.000 | -0.3224 | -6.24 | 0.000 |
| Negative opinion of the quality of NHS care services | -0.1724 | -3.48 | 0.001 | -0.1734 | -3.50 | 0.001 |
| Negative opinion of existing LTC services | -0.0710 | -1.52 | 0.128 | -0.0718 | -1.54 | 0.124 |
| State should pay basic LTC services to all | -0.4730 | -9.20 | 0.000 | -0.4745 | -9.23 | 0.000 |
| State should pay basic LTC services only to the poor | -0.1563 | -2.84 | 0.005 | -0.1569 | -2.85 | 0.004 |
| *IV parameters* | | | | | | |
| No insurance | unidentified | | | unidentified | | |
| Insurance | 0.5487 | 14.545 | 0.000 | 0.566 | 14.535 | 0.000 |
| *Diagnostic statistics and tests* | *Value* | | | *Value* | | |
| Log likelihood function | -11117.7 | | | -11102.1 | | |
| Pseudo R-squared | 0.156 | | | 0.158 | | |
| Number of observations | 11760 (1176) | | | 11760 (1176) | | |

Different theoretical arguments can be used to explain the relation between demand for a more ample coverage and age. The most intuitive one is that elderly people are more willing to contribute to the programme since they are more likely to be in need of LTC in the near future. Similarly to what happens with traditional health care insurance, a larger asset variance in expected expenditures for care should produce this effect. Our results contend the empirical relevance of this effect, however, in line with Becker and Zweifel (2004) who study hypothetical demand for health insurance in Switzerland in a stated preferences framework. This result can be rationalised within the framework of theoretical analyses, such as Meier's (1999), who have stressed the importance of uncertainty in modifying incentives to demand LTC coverage with age. In particular, greater uncertainty over the probability of becoming disabled should favour early purchase of LTC insurance, whereas uncertainty over future costs of disability favour late purchase of insurance. Our results suggest that the first issue seems to be more relevant. This is consistent with the structure of the experiment, where costs of disability were explicitly reported, and with the nature of LTC expenditure, which is much more standardized than general health expenditure.

An additional possible explanation is that younger people also fear that current welfare programs might no longer be financially sustainable once they reach dependency age. Ageing of the population and increasing restraints on the public budget may limit the provision of an adequate level of coverage for future generations, who therefore are more interested in extending current programs. Clearly, what emerges from our results is that the lack of coverage in real markets cannot simply be attributed to the myopic attitudes of young generations.

Individuals with poor self rated health state are more likely to opt for the status quo. A variety of explanations may be viable. In particular, people who suffer of a generic bad health state could presume to qualify for free social care already under current legislation. Interestingly, however, whilst generic bad health conditions do not increase demand for coverage, chronic conditions and hospitalisation in the year prior to the survey both have a positive influence on the probability of opting for more extensive coverage. People who have experienced chronic illnesses are probably more aware of the high (monetary and non-monetary) burden that individuals are currently obliged to assume. They probably already receive some kind of help (either informal or publicly provided) or they perceive the risk of needing assistance in the near future as particularly high and in both cases the benefits from greater coverage are highly evaluated.

Another group of variables included in the regression refers to respondent's educational and employment status. For education, the reference category represents non educated respondents, and all the coefficients are statistically significant with absolute value increasing with the level of education. This positive correlation between education and propensity to cover is probably due to a higher awareness of the difficulties to face the actual burden (and the expected increase) of the costs related to informal and formal care. The result is consistent with empirical evidence provided by the revealed preferences literature that studies the demand for both LTC (Mellor, 2001) and supplementary medical insurance (Besley, Hall and Preston, 1999) where most educated households are more likely to purchase coverage.

On the contrary, individuals' work status plays a minor role in the decision process. White and blue collar workers do not reveal any significant difference with respect to the self employed, assumed as base case. Such a result is not totally surprising if one considers that the argument of greater opportunity costs of illness for self employed individuals is weakened for the kind of coverage we are considering here since, differently from standard health insurance policies, LTC coverage acquires increasing importance with ageing, and many individuals will experience disabilities after retiring. Still, the coefficients for the retired and non-occupied condition (though referring to a limited number of respondents) are significant, with the former being more likely to choose the status quo.

We have also studied the effects of a few "opinion variables". A negative opinion on the quality of care currently provided by the National Health Service favours an extension of coverage. Individuals do not seem to respond to unsatisfactory quality of public health care by relying on out-of-pocket expenditures, but to look for additional financial support for ancillary programs such as the one proposed here concerning LTC. Our estimated model also contains an indicator of respondent's opinion on existing LTC services, which does not show any significant role.

Finally, being a subscriber to a private health insurance policy has a positive impact on the likelihood of being willing to contribute to LTC coverage. The result is consistent with our expectations. On the one hand, policy holders are expected to be more risk averse, and to perceive the insurance mechanism as an effective tool for facing health related risks. On the other hand, familiarity with insurance products reduces perceived transaction costs with regard to making use of a (new) policy.

## 4.1 Evaluation of attributes and welfare analysis:

Drawing on the results discussed above, we have estimated the marginal price or WTP where coverage against LTC risk is itroduced. The issue of deriving welfare measures from discrete choice experiments has been recently and widely debated in the health economics literature (Lancsar and Savage, 2004; Ryan, 2004; Santos Silva, 2004). Following Ryan's (2004) classification, the subject of interest here is the estimation of welfare effects related to a "state-of-the-world-model", i.e. a situation where the kind of good or service will be obtained by an individual is known with certainty (see also Louvière *et al.*, 2000, p. 337). In this case, the welfare measure, namely the compensating variation, for a change in the characteristics of an available alternative is:

$$WTP = -\frac{1}{\beta_p}\left(V_0^h - V_1^h\right), \tag{17}$$

where the subscripts (0, 1) define indirect utility functions before and after the policy change, and $\beta_p$ is an approximation of the inverse of marginal utility of income, which is recovered from the estimated coefficient of the variable expressed in monetary terms. If only one attribute is changing, then we obtain an "implicit price", or marginal WTP. This expression also represents an appropriate measure for those cases where a quality variation applies to all the alternatives of the choice set (Haab and McConnell, 2002).

As long as WTP is determined as a difference between utility functions, it follows that in our model only the attributes determine the welfare measure, and that the utility index at the status quo can be set to zero.

Implicit prices for each attribute are reported in the upper part of Table 4, and are derived from the parameter estimates reported in Table 2 and 3 by dividing the estimated coefficients of the non monetary attributes by the negative of the cost coefficient. Confidence intervals (at 95%) have been computed with the Krinsky-Robb procedure in order to assess the robustness of the results.

Table 4: *Implicit prices of the attributes and mean WTP for hypothetical schemes*

| | *Multinomial estimations* | | | | *NL estimations* | | | |
|---|---|---|---|---|---|---|---|---|
| | *MN logit* | | *MN probit* | | *Base Model* | | *Model with interactions* | |
| *Attributes of the coverage programme* | *Implicit prices in Euros* | *K-R 95% confiden interval* | *Implicit prices in euro* | *K-R 95% confiden interval* | *Implicit prices in Euros* | *K-R 95% confiden interval* | *implicit prices in Euros* | *K-R 95% confiden interval* |
| 1% Degree of coverage | 11.04 | *10.03 12.18* | 12.12 | *10.08 14.73* | 11.03 | *8.49 12.80* | 10.50 | *8.49 12.8* |
| Option to cover residential costs | 345.0 | *305.1 387.7* | 345.7 | *280.9 425.0* | 314.9 | *256.1- 384.1* | 351.7 | *284.9 430.4* |
| Option to cover residential costs if 30% coverage | | | | | | | 155.4 | 51.8 267.9 |
| Difference between private insurance and public cover | 183.6 | *149.2 219.2* | 197.2 | *152.2 249.5* | 178.1 | *137.2- 226.1* | 220.3 | *168.5 281.1* |
| Difference between private and public cover if 100% cover | | | | | | | 122.7 | *35.9 213.3* |
| ASC (*0 for status quo*) | -864.0 | *-775.6 - 961.0* | -756.2 | *-521.2 - 924.2* | -1066.8 | *-619.4 - 1541.0* | -1027.1 | *-589.5 - 1505.9* |
| | *WTP for representative scenarios* | | | | | | | |
| Scenario | ***Mean WTP*** | *K-R 95% confiden interval* | ***Mean WTP*** | *K-R 95% confiden interval* | ***Mean WTP*** | *K-R 95% confiden interval* | ***Mean WTP*** | *K-R 95% confiden interval* |
| 100% of coverage no option for residential costs — *Private cover* | 240.3 | *143.9 340.6* | 502 | *270.6 759.8* | 36.0 | *-411.2 497.8* | 19.4 | *-440.8 481.6* |
| 100% of coverage no option for residential costs — *Public cover* | 423.8 | *320.1 531.9* | 699 | *460.4 971.0* | 214.1 | *-233.8 677.6* | 142.1 | *-322.7 615.2* |
| 65% of coverage with option for residential costs — *Private cover* | 198.8 | *113.5 289.6* | 423.5 | *217.9 646.9* | -35.2 | *-470.9 411.1* | 4.8 | *-445.5 451.3* |
| 65% of coverage with option for residential costs — *Public cover* | 382.4 | *289.4 481.1* | 620.7 | *406.1 860.1* | 142.9 | *-298.1 589.7* | 225.1 | *-222.2 677.2* |
| 100% of coverage with option for residential costs — *Private cover* | 585.3 | *478.1 700.7* | 847.9 | *597.1 1135.0* | 350.8 | *-101.1 813.6* | 371.1 | *-85.6 847.6* |
| 100% of coverage with option for residential costs — *Public cover* | 768.9 | *653.9 893.4* | 1045.1 | *785.6 1353.8* | 528.9 | *70.5 999.0* | 493.7 | *27.6 985.7* |

In the "main effects" MNL and NL models, the estimated marginal WTP amounts to 11.0 Euros per 1% increase in coverage (i.e. reduction of the copayment rate). The WTP from the NL model is only slightly affected by the inclusion of the two significant interaction terms. The MNP specification yields the highest value (12.1 Euros).

A remarkable result is the high value attached to the extension of coverage to residential care expenditures (315 and 352 euros in the two NL specifications; 345 euros in the two multinomial models), which testifies a strong concern for the risk of being forced to leave one's own domicile. Finally, the differential utility for the public solution, amounts to 178 Euros of additional WTP with the NL base model. For this attribute, the difference between main effects and end point design estimates is quite relevant. Public coverage becomes less important (123 euros) when all LTC expenditures are insured.

We have used equation (17) to get an estimate of the overall mean WTP. Mean WTP has to be estimated taking into account the negative value of the alternative specific constant in order to capture the utility loss that would occur in cases where there is compulsory introduction of the programmes for those respondents who would opt for the 'status quo'. The results in the bottom part of Table 4 refer to a few hypothetical schemes and show that the difference between the multinomial and the NL estimates is quite large. This is mainly due to the different ASC value, since with the NL model we have considered a series of individual specific regressors which partly explain the variance of the "status quo" choice. Differences in ASC estimates are also responsible for the much larger confidence intervals for the mean WTP estimates in the NL specification, due to the lower significance level.

WTP estimates can serve for exploring not only the possibilities of expanding private markets for LTC insurance but also the political sustainability of implementing new public programs, and represent a useful benchmark for assessing the potential support of such policy innovations. Since our analysis is mainly focused on demand, a precise estimate of the actual costs of providing additional LTC coverage for public and private insurers is beyond the scope of the paper. However, some indications that broadly reflect the supply side conditions in the Italian context can be collected, either from existing studies or from a direct inspection of the (narrow) existing private markets, and can be compared to mean WTPs obtained from our experiment, in order to make some conjectures on the extent to which supply can match demand for coverage.

Existing supply-side studies in Italy are micro-simulation exercises carried out on a regional basis mainly to provide policymakers with some broad indications on the approximate tax price for expanding public guarantees in the LTC area, but which usually do not aim to develop accurate forecasts for future financial needs.They consequently suffer from some ambiguities in the exact nature of the claims citizen will be entitled to, and on the kind of services involved. More in general, the coverage scenarios considered are not fully comparable with those we proposed. Ahead of these limitations, Lemmi and Sciclone (2003) have estimated an average tax price increase of 392 Euros per household for implementing a comprehensive coverage program for LTC in Tuscany (a region which has many similarities with Emilia Romagna in terms of socio-demographic characteristics and quality and organisation of welfare services). The authors hypothesise two alternative financing schemes both based on personal income taxation. One with a uniform increase of 2.4% in the tax rate, and a more progressive one where the tax rate increases from 2.2% up to 3.3%. A similar policy change is explored by Coda Moscarola (2003) for a different Italian region, Piedmont. She obtains that if all citizens above 24 years of age were asked to contribute, the personal fiscal contribution would range from 383 to 340 Euros according to whether services are more oriented towards residential or home care.

As for the information that can be recovered from private markets, one can infer an estimate of costs of coverage by referring to the present prices of the LTC policies. Such products have been introduced in Italy only recently, and the number of policies sold is still very limited. Prices are usually set after collecting personal information on the potential subscriber through a questionnaire but, at least for some insurance companies, it is possible to gather information on their current average prices, conditional on observable individual characteristics such as age, gender and (sometimes) province of residence. Again, any comparison between such prices and our estimates must be taken very cautiously. First, the benefits policies will afford vary according to severity, whereas we were forced to limit our scenario to a single health status for minimising cognitive difficulties of the respondents. Moreover, real LTC policies may contain additional clauses not considered in our simplified hypothetical package. Finally, the way benefits are provided may differ substantially between policies (e.g., some companies concentrate relatively more on home care support, whilst others are more oriented to cover residential care needs; some may provide only cash benefits, while others directly supply care through a network of selectively contracted providers). In spite of all this, price information for some roughly comparable packages can be collected, in particular by referring to insurance policies that merely ensure a predefined amount of money when disability occurs (indemnity policy scheme). Table 5 reports actual prices of individual LTC policies that cover different monetary amounts for two insurance companies.

Table 5: *A few yearly premiums from the Italian LTC insurance market (Nov. 2005)*

| | ***Yearly premium in Euros*** | | |
|---|---|---|---|
| | *Low Coverage (€ 516)* | *Medium Coverage (€ 1033)* | *High Coverage (€ 1550)* |
| **Male** | | | |
| Insurance Company A | 194 | 388 | 583 |
| Insurance Company B | 241 | 474 | 693 |
| **Female** | | | |
| Insurance Company A | 258 | 516 | 774 |
| Insurance Company B | 338 | 668 | 984 |

The scenarios considered in the bottom part of Table 4 can be compared to the values of Table 5. What can be inferred is that the relationship between WTP and indemnity insured resulting from our experiment is much steeper than the relationship emerging from market products. It turns out that a match between demand and supply is in theory possible only for high coverage levels.

In our analysis, we found that on average people are willing to pay more if programmes are organised by public authorities. For individuals of average age, WTP for private insurance is substantially below what subscribers are asked to pay in the market. Hence, expected benefits of private coverage do not seem to be evaluated great enough by the average consumer to compensate for insurance prices. This contributes to explaining the well-known empirical evidence concerning the small size of LTC insurance markets.

On the other hand, preferences for public coverage are such that estimated WTP does not differ excessively from the amount that seems necessary for extending coverage to an extent that regional health planners are actually considering. Yet, nearly 25% of the respondents always prefer the status quo; and the probability of their being willing to extend coverage increases with income. This may raise some relevant opposition to the programme in particular in low income groups, who have probably more urgent priorities which could justify an increase of the tax burden.

Interestingly, the difference between cost for coverage and WTP is not homogeneously distributed across the population. For instance, in market policies males are required to pay a lower premium than females, presumably because of a lower life expectancy. At the same time, in our sample they are more willing to choose greater coverage. A similar result holds for age, with younger individuals having cheaper access to private policies in the market but displaying at the same time a larger propensity to cover in our hypothetical exercise. This suggests that if insurance companies want to enlarge the market for LTC policies, they should target young well-off males in particular, because the gap between expected cost of coverage and WTP is by far lower than in other socio-economic groups. Such cream skimming behaviour would leave unsolved the social problem of ensuring coverage to the other socio-demographic groups.

## 5 Concluding remarks

This paper has analysed the results of a discrete choice experiment carried out on a representative sample of the population of the Italian region Emilia-Romagna. The experiment was aimed at inferring the characteristics of potential demand for LTC risk insurance and eliciting the WTP for some policy prospects. An analysis based on a stated preference approach is particularly useful for policy decisions when there is scarcity of information from real data, and for comparing different institutional scenarios.

A basic hypothetical scenario was varied according to the levels of four main attributes which defined the LTC coverage: the yearly cost of the insurance premium, the payment scheme (voluntary subscription to a private insurer vs. compulsory personal income taxation), the option right to extend coverage to residential care and the co-payment rate. These attributes were all highly statistically significant, confirming that the characteristics we considered are all influential and appropriate for the individuals.

From a methodological viewpoint, we developed the idea to provide the "opting out" option with some economic structure, something often disregarded in the health economic literature. Many variables considered in the nested logit specification proposed have highly statistically significant effects, and indications concerning whether extending current LTC coverage or not are probably as economically relevant as WTP measures for LTC cover. Moreover, modelling the demand for insurance in a choice experiment setting allows for investigating pure demand effects, which are difficult to identify because markets can be crowded out by public intervention and, even when private markets exists, are often hindered by the interference of supply side constraints, due to the incentives for private insurers to select risks (e.g. Propper, 1993).

An important indication derived from the empirical analysis is that preferences for extending coverage for LTC risk are heterogeneous. Around one fourth of the sample opts for the status quo, whereas the remaining fraction prefer greater levels of coverage than what is currently ensured by the public sector. Moreover, a systematic influence of many socio-demographic characteristics on the decision of whether or not to extend coverage has been detected. When individual preferences are very heterogeneous, private rather than public oriented solutions tend to be preferred, since the former better preserve consumer's sovereignty (e.g. Becker and Zweifel, 2004). However, against this general backdrop, two aspects emerge in our analysis. First, the dimension over which preferences are shown to differ most is the extension of coverage, rather than the structure of the insurance package. According to our modelling, what could mismatch preferences under public coverage is the fact that compulsory payment is asked of some groups even if they would not be willing to pay. However, since empirical analysis enables us to distinguish those groups which benefit less from additional insurance, public authorities could minimise welfare losses by introducing exemption thresholds according to some observable characteristics such as income or age. In this way, the non negligible political opposition to the reform would be probably attenuated, and at the same time distortions in individual choice would be reduced. A second important aspect relates to the fact that, in some areas of intervention, preferences depend on the nature of the organisation providing a particular services. This is likely to be the case for elderly care. In our sample, respondents on average are ready to pay an extra-premium for having the same coverage publicly rather than privately provided. This further attenuates the fears of possible welfare losses due to public provision. Overall, our results provide a rationale for measures such as threshold exemptions, which are often also observed in practice, and are here motivated on mere efficiency rather than equity grounds, as typically happens.

The welfare estimations derived from the regression results display a fairly high mean WTP, with a value of 10-11 euros per percentage point of the copayment rate. This indication leads to substantial WTP when the copayment is very low which, at least for some socio-demographic groups, does not differ excessively from insurance policies actually sold in the market. As a note of caution, we must point out that the application of stated preference techniques to insurance markets is a difficult task, and that the use of the results must be prudent. For example, we still find it unsatisfactory that a large part of overall WTP is not captured by the attributes of our hypothetical coverage scheme, despite their high significance level, or by the individual-specific variables included in the nested logit regressions, but by the alternative specific constant. Apparently, preferences over a very complex issue such as the one analysed, are by far more articulated than it was possible to model with our analysis. Further improvement both at the methodological and economic level are therefore necessary in order to provide more in-depth insights on the subject of LTC insurance demand.

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