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Contracting out refuse collection

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Abstract. In this paper we seek an explanation for the reservations of local authorities towards contracting out. Although empirical evidence suggests that contracting out results in a significant cost decrease, a majority of Dutch municipalities provides for waste collection services themselves. Based on theoretical insights we model the choice between private, public, in-house, and out-house refuse collection. The models are estimated using a database comprising nearly all Dutch municipalities. We find evidence that the number of inhabitants, the transfer by central government, and interest group arguments are important explanations. Interestingly, ideology seems to play a minor role.

Compared to earlier studies we estimate more general models. Although the same qualitative results are found for parametric and semiparametric models, we find strong statistical evidence that a parametric specification is far too inflexible. Differences between the parametric and the semiparametric marginal effects are substantial. Thus, more attention is needed for the implications of model specification.

Key words: Refuse collection, institutional choice, ideology, interest groups, semiparametric estimation

1. Introduction

There seems to be evidence that contracting out government services saves taxpayers money, and sometimes a lot of money, compared to public provision. In an overview, Domberger and Jensen (1997) show that contracting out a broad field of government services might result in cost savings in the order of 20% without sacrificing the quality of services provided.

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Also Tang (1997), in a critical assessment of several studies, comes to the conclusion that the private sector is found to be more efficient in refuse collection, fire protection, cleaning services, and capital intensive waste-water treatment, while in sectors as water supply and railways the results are more mixed.

Especially, the cost savings of private refuse collection have been discussed at length in the literature. Kitchen (1976) estimates a cost decrease of Canadian \$2.23 per capita when private firms collect household waste. Stevens (1978) arrives at a cost decrease between 7% and 30% due to contracting out for the USA, where the magnitude of the effect depends on the size of the municipality. Based on UK-data Domberger et al. (1986) published a study on the effects of contracting out household refuse collection in the United Kingdom. They concluded that there are cost savings of 22% for contracting out to private companies. Szymanski and Wilkins (1993) and Szymanski (1996) have confirmed these results, based on an extension (in years) of this database. Dijkgraaf and Gradus (1997) show similar cost savings between 15% and 20% for the Netherlands, in case Dutch municipalities are contracting out refuse collection. Moreover, Ohlsson (1998) reports almost the same estimations for Sweden. Recently, Bosch, Predaja and Suarez-Pandiello (2000) presented Spanish data for 73 municipalities in Catalonia. They pointed out that the framework for which the service is provided is more relevant than the public private dichotomy. In a recent contribution Reeves and Barrow (2000) pointed out cost savings of around 45% in Ireland.

Although the practice of contracting out refuse collection has become more popular, it is still less common than in-house provision. In the United Kingdom only 30% of the contracts for refuse collection is placed out-house (see Szymanski (1996)). According to Reeves and Barrow (2000), in Ireland in 39% of the studied cases private providers were contracted to provide refuse collection. In the Netherlands 40% of the municipalities use private collectors for refuse. However, due to the fact that private collectors are especially active in small villages, only 20% of total tonnage is in private hands (see Dijkgraaf and Gradus (1997)). Only Ohlsson (1998) finds for the Swedish case that private provision is slightly more common than public provision.

Furthermore, a recent study by López-de-Silanes et al. (1997) shows the reservations of local authorities towards contracting out. Based on data in 1987 and 1992 for 3042 counties for twelve services like water supply, landfills, libraries etc. only 25% of the services in 1987 and 35% in 1992 had been placed out-house. Moreover, in this article a nice empirical investigation of the mode of providing government services is given, where three leading theories (namely efficiency, political patronage, and ideology) are investigated. The evidence presented in this article indicates that clean government laws and state laws restricting county spending encourage privatisation, whereas strong public unions discourage it. This suggests an important role played by political patronage and taxpayer resistance to government spending in the privatisation decision.

In this article, we examine for the Netherlands the determinants of the provision mode of refuse collection. Data are available for 540 (i.e., almost all) Dutch municipalities. We find evidence for political patronage and the wealth of the local government as a ground for contracting out, but also the possible efficiency gain of contracting out plays a role. Moreover, we extend the existing literature by investigating more general specifications. Especially, the usually applied logit model seems too restrictive. Formal tests strongly reject the appropriateness of the logistic probability transformation. As alternative we use a semiparametric single index modelling approach, based on Ichimura (1993), where the probability transformation is left unrestricted. We find that the semiparametric single indices are comparable to the parametric analogues, but the probability transformations are quite different, implying that the logit specification might yield misleading predictions, particularly, when considering marginal effects.

The remainder of this paper is organised as follows. In Section 2 we discuss the relevant theoretical issues. In Section 3 we describe the data we use. Section 4 contains the estimation results based on logit. In Section 5 we investigate the robustness of these results, by testing the logit specification, and by using a semiparametric alternative, based on Ichimura (1993). Section 6 concludes.

2. Theoretical issues

Before we specify the data and the empirical results, it is worthwhile to discuss some theoretical issues concerning the contracting out decision (see also López-de-Silanes et al. (1997), and Tang (1997)). As mentioned in the introduction, Dijkgraaf and Gradus (1977) show that Dutch municipalities might achieve cost savings between 15% and 20% in case of contracting out refuse collection. With lower service costs, one would expect that municipalities favour private collection. Indeed, 40% of the Dutch municipalities chose for the option to collect waste by a private firm. The question arises: why did the other 60% not choose this option as well?

Hart et al. (1997) argue that private contractors might fail to pursue goals that politicians want to attain. Especially, in circumstances such as health care and prisons, where politicians cannot write a complete contract that specifies exactly what contractors are supposed to do in all circumstances, it may not be straightforward to contract out. The logic suggests some potential efficiency benefits of in-house government services to ensure quality. However, it is not clear how important such benefits are for refuse collection. Hart et al. (1997, p. 1154) argue that in the case of refuse collection the damage to quality can be offset by a good contract, so that "private provision is superior". Nevertheless, according to a Dutch inquiry, such elements are still available and some municipalities put forward that quality is the reason for in-house provision (see NG *magazine* (1998)). A prediction following from this kind of reasoning is that the wealth of local government decreases the likelihood of contracting out. A poorer government is less likely to care about quality and is more interested in cost savings.

Related to these wealth arguments are the so-called output arguments. Some empirical insights suggests a linear relation between the cost of service and output (number of inhabitants, pick up points etc., see, for example, Domberger et al. (1986)). However, especially for small municipalities this may not be true. Kitchen (1976) finds that the maximum scale in refuse collection occurs in cities of about 324,000 inhabitants. Stevens (1978) divides the sample into several subsamples. For small municipalities there is less evidence for this linear relation. Therefore, she finds increasing returns to scale, if the city population is less than fifty thousand and constant returns to scale if the city population is larger than fifty thousand. A prediction following from this kind of reasoning is that the number of inhabitants decreases the likelihood of contracting out. However, this relation may not be linear. Above a certain level there is less evidence that private waste collectors have more opportunities to combine the collection of different municipalities and thus to use scale effects as a cost decreasing mechanism.

An alternative view of the contracting out decision focuses on public choice theory (see Buchanan (1987)). This approach explains social behaviour as the product of free choices of individuals. Self-interested politicians, bureaucrats and unions have a stake in in-house provision as they can use it as a status-enhancing feature. López-de-Silanes et al. (1997) argue that in the United States the main political factor favouring in-house provision seems to be the public employee unions. Moreover, the role of unions becomes more important and, therefore, in-house provision becomes more beneficiary if unemployment in a municipality is high.

The third theory stresses the importance of voter ideology. To evaluate this view, one should take into account voting patterns in different municipalities. Hereby, it is assumed that the contracting out decision is simultaneously determined by the degree of voters' anti-government sentiment. This laissez-faire sentiment is most visible in right-wing parties.

Finally, it is possible that the privatization decision in a particular municipality is related to what happens in other municipalities. For instance, Bivand and Szymanski (2000) find evidence for the UK that in the period before Compulsory Competitive Tendering (CCT) costs were spatially correlated across authorities, while following CCT this spatial correlation disappeared. To account for this effect, Bivand and Szymanski suggest that before CCT most local authorities evaluated the service costs by comparison with their local neighbours. Municipalities with a higher than average cost compared with the neighbours would choose the option of privatization. In addition, the decision of contiguous municipalities might affect the decision of a municipality via scale economy, especially when the municipality under consideration is small. Alternatively, one could argue that municipalities might take into account the decisions in some kind of reference group of municipalities, where the reference group consists of municipalities which are, for instance, comparable in size or in number of inhabitants.

However, contrary to the first three points, this fourth issue, interdependence between municipalities is much harder to quantify. Without knowledge of which municipalities influence which municipalities, the researcher will have to model such interdependencies him- or herself by modelling reference groups. However, as argued by Manski (1993) in the context of a linear demand equation for consumers with interdependencies between consumers, it is impossible to infer unknown reference groups on the basis of observed behaviour: an informed specification of reference groups is a necessary prelude to an analysis of interdependent behaviour. As such information is not available for our case estimation of the effects of interdependencies is not possible.[1]

[1] Moreover, since our model is of a binary choice type, we would also have to deal with the problem of "coherency", when modelling interdependencies, see, for example Schmidt (1981) or Gourieroux et al. (1980): the interdependency should be of a recursive type (one municipality may influence the other, but then not the other way around), since otherwise the model is not coherent, i.e., probabilities do not sum to one.

Table 1. Descriptive statistics database

Variables	Average	Maximum	Minimum	St. dev.
Private provision (%)	42	100	0	49
In-house provision (%)	28	100	0	45
Inhabitants (×1000)	26	722	1	45
Inhabitants per hectare	6	63	0	8
Transfer from central government per inhab. (euro)	442	1727	118	113
Income per inhabitant (1000 Euro)	9	13	6	1
Unemployed per 100 inhabitants	3	6	1	1
Local civil servants per 100 inhabitants	11	16	8	3
Conservative Liberals (%)	16	52	0	9
Social Democrats (%)	16	49	0	9
Progressive Liberals (%)	8	34	0	7
Orthodox Protestants (%)	6	67	0	10
Green Left (%)	4	34	0	6
Extreme Right (%)	0	11	0	2
Local parties (%)	25	100	0	20

3. Data

To test the theories about contracting out, a database is constructed with data on the different institutional forms of waste collection and variables representing the theories. The data on the different institutional forms is based on a 1998 census of the Dutch Association for Refuse and Cleansing Management (NVRD). Moreover, municipalities' characteristics are available from Statistics Netherlands (CBS). For 540 of the Dutch municipalities (96% of all municipalities) figures are available, see Table 1.

Institutional forms

In general, three modes of provision are used in this dataset. The first mode is provision by a private firm (42%). The second and third mode are both by public ownership but differ with respect to autonomy of the collection service. The second mode occurs when municipalities collect the waste of their own citizens (28%). The waste collection service is in this case under direct control of the municipality council. The third mode occurs when another municipality or an external public organisation (30%) collects the waste, so that the municipality council has less direct control on the waste collection service.

Output variables

To check for the output arguments the number of inhabitants and population density (number of inhabitants per hectare) are included in the empirical setting. On average a Dutch municipality has 26 thousand inhabitants, while the largest city (Amsterdam) has 722 thousand inhabitants and the smallest municipality only 1 thousand. To check for scale economy the number of inhabitants squared is included as well. Moreover, the population density shows a high variation between municipalities, indicating that the transport distance between individual pick-up points varies.

Wealth variables

The theory about the influence of wealth on contracting out suggests that budget constraints influence the trade-off between efficiency and social arguments. Hard budget constraints increase the likelihood of privatisation. In the Netherlands the income of local government depends almost totally on the transfers by the central government. The freedom of Dutch municipalities to collect their own taxes is quite restricted. Therefore, we include as an explaining variable the transfer from central to local government per inhabitant. As the trade-off between efficiency and social arguments depends on the social characteristics of the inhabitants we include the average personal income in a municipality as a wealth variable as well. The hypothesis is that a municipality will weigh cost savings more if the inhabitants are poor.

Interest group variables

In the López-de-Silanes et al.-study interest group variables are included for the number of public employee's or union membership and for the opportunity to purchase supplies from political allies (the so-called clean government variables). However, for the Netherlands clean government laws are dictated at a national level and, therefore, these data cannot be included. No data are available for the number of public employee's in a municipality. However, these data are available at a regional level and are, therefore, included.[2] Similar to López-de-Silanes et al., it is possible to include labour-market conditions as an approximation of interest group variables. In general, we should expect that in-house provision becomes more beneficiary if unemployment in a municipality is high. Therefore, the unemployment level is included in our estimations.

Political variables

We include the fractions of the following parties, based on the local elections of May 1994[3]: green left, social democrats, conservative liberals, progressive liberals, orthodox Protestants, extreme right and local parties.[4] In the estimations the Christian democrats, who are in the middle of the political spectrum, are excluded.[5]

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[2] There are twelve provinces or regions in the Netherlands.

[3] There were new elections in May 1998.

[4] Green left: Groen Links + SP, social democrats: PvdA, conservative liberals: VVD, progressive liberals: D66, Christian democrats: CDA, orthodox Protestant: SGP + RPF + GPV, extreme right: CD and local parties: other parties. Combination of the parties is tested using a Log Likelihood test.

[5] In addition, we looked at municipality-level voting in the 1994-election for Parliament as alternative indicator of the electorate's ideological orientation. However, local elections seem to be the best means of predicting the probability of private contracting.

4. Estimation results: logit

We start our estimations with a standard logit analysis for two models.[6] In the first model, the choice between public and private provisions is estimated as dependent on a number of explaining variables. In the second model the choice between in-house and out-house provision is the dependent variable. In both models, all explaining variables are initially the same.[7] Thus, the basic model is:

$$P(Dep = 1 \mid x) = \Lambda(\beta^T x),$$

where:

Dep: Dependent variable,
model 1: dummy with value 1 for municipalities with no private collection;
model 2: dummy with value 1 for municipalities with collection in-house;

and where x contains the following explanatory variables (next to a constant term):

Inhabitants	Number of inhabitants (*10000);
Funds	Transfers from central government (Euro per inhabitant);
Income	Personal income (Euro per inhabitant);
Civil servants	Number of civil servants (per 100 inhabitants);
Unemployment	Number of persons with an unemployment benefit (per 100 inhabitants);
Conservative Liberals	percentage of total votes in a municipality;
Orthodox Protestants	percentage of total votes in a municipality;
Social Democrats	percentage of total votes in a municipality;
Progressive Liberals	percentage of total votes in a municipality;
Green Left	percentage of total votes in a municipality;
Extreme Right	percentage of total votes in a municipality;
Local Parties	percentage of total votes in a municipality.

To account for sufficient flexibility in terms of the number of inhabitants, we also decided to include the number of inhabitants squared (/1000). The parameter vector β contains the unknown parameters, and Λ represents the logit-transformation.

Results are given in Table 2. This table contains the estimated parameters, together with the estimated standard errors.

First, we discuss the **no-private provision** case.

Output variables

It shows that scale effects are present. The estimated second order polynomial in terms of inhabitants is increasing up to its maximum at around

[6] The probit and the OLS results are extremely similar.
[7] An interesting extension would be to include the previous state of the dependent variable as an explanatory variable. However, such data are not available.

Table 2. Estimation results logit and Ichimura model (estimated standard errors in brackets)

Variables	No-private collection (logit)	No-private collection (Ichimura)	In-house collection (logit)	In-house collection (Ichimura)
Constant	−1.72	–	−4.51	–
	(1.88)	–	(2.18)	–
Inhabitants	0.26	0.26	0.18	0.18
	(0.11)	–	(0.08)	–
Inhabitants squared	−4.16	−3.94	−2.45	−2.57
	(1.73)	(0.30)	(1.13)	(0.38)
Population density	0.08	0.07	0.03	0.05
	(0.03)	(0.03)	(0.02)	(0.01)
Fund	1.38	1.11	1.36	1.34
	(0.66)	(0.55)	(0.73)	(0.33)
Income	−1.61	−1.24	−0.67	−1.17
	(0.93)	(0.73)	(1.05)	(0.43)
Unemployment	0.02	0.02	0.10	0.04
	(0.02)	(0.02)	(0.03)	(0.01)
Civil servants	0.27	0.19	0.10	0.04
	(0.05)	(0.05)	(0.05)	(0.02)
Conservative Liberals (%)	−0.004	−0.01	0.0008	0.009
	(0.02)	(0.02)	(0.02)	(0.01)
Social Democrats (%)	−0.0003	−0.02	−0.04	−0.03
	(0.02)	(0.01)	(0.02)	(0.01)
Progressive Liberals (%)	0.008	−0.02	0.03	0.01
	(0.02)	(0.01)	(0.02)	(0.01)
Orthodox Protestants (%)	0.03	0.01	0.02	0.009
	(0.01)	(0.01)	(0.02)	(0.01)
Green Left (%)	−0.02	−0.04	−0.01	−0.01
	(0.02)	(0.02)	(0.02)	(0.01)
Extreme Right (%)	−0.05	−0.08	−0.18	−0.29
	(0.13)	(0.06)	(0.08)	(0.04)
Local party (%)	−0.02	−0.03	−0.03	−0.008
	(0.01)	(0.01)	(0.01)	(0.01)
Log likelihood	−281.52		−256.40	

312,500 inhabitants, so that with an increasing number of inhabitants (up to this maximum) the probability of public provision (i.e., no private provision) increases.[8] The occurrence of scale effects makes public provision more likely. Furthermore, if the number of inhabitants per hectare increases the probability of public provision increases. Again scale effects are present.

Wealth variables

As we expected, more transfers by the central government favours public provision, because less emphasis has to be given to cost savings investigations. Contrary to our prior, a higher income level in a municipality lowers the probability of public provision. However, the estimated coefficient is not significant.

[8] In the Netherlands only three cities have more inhabitants than this maximum, namely The Hague, Rotterdam, and Amsterdam.

Interest group variables

Interesting are the results with respect to the interest group variables. The data give evidence for the prior that the number of public employees raises the probability of public provision. Also the number of unemployed persons raises the probability of public provision, although the coefficient estimate is not significant.

Political variables

The results with respect to the political variables are much weaker.[9] Only local parties are against public provision in a significant way (compared to the Christian Democrats, who are the reference group). Probably, this can be explained by the anti-government sentiment by some of these local parties. From the other parties only the Orthodox Protestants are in favour of public provision in a significant way. This can probably be explained by the reserved attitude towards the role of market forces in these parties.

For **in-house provision** the over-all results are in line with no-private provision. The top of the polynomial in terms of inhabitants is now at around 367,000 inhabitants: thus, if the number of habitants increases (up to this maximum) then the probability of in-house provision increases. The effect of inhabitants per hectare, however, becomes insignificant. This also applies to the effect of transfers by the central government. For the number of unemployed persons the effect of in-house provision seems somewhat stronger and quite significant. The effect of income per inhabitant remains contrary to our prior, but the estimated coefficient is again not significant. The effect of the number of public employees is again positive and significant. This seems to be in line with the theory that interest group considerations are an important obstacle to out-house provision. In addition, the results for political variables are also here suggestive. The attitude of the social democrats and extreme right towards in-house provision turns out to be significant, whereas the effect of the other parties is insignificant (compared to the Christian Democrats).

5. Robustness of results

The basic logit model presented in the previous section requires strong distributional assumptions to be valid. In particular, the assumption that the probability transformation is given by the logistic probability distribution Λ may be questioned. To investigate the validity of this assumption we tested it against a more general specification as proposed by Ruud (1984), and as used

[9] The insignificance of political variables may be sensitive to specification of these variables. Therefore, we experimented with a "left/right" variable. This "left/right" variable is constructed as follows: 8*Green Left + 7*Social Democrats + 6*Progressive Liberals + 5*Christen Democrats + 4*Local Parties + 3*Orthodox Protestant + 2*Conservative Liberals + 1*Extreme Right. A log-likelihood test was used to investigate whether the model with individual parties is preferred. We obtain as test result for the no-privatisation case a value of 25.86 and for the in-house modelling a value of 27.52; since the test is asymptotically chi-squared-distributed with 6 degrees of freedom, the tests lead to rejection at the 1%-level, suggesting that the specification including different parties is preferred. Furthermore, the left/right variable is not significant at 10%.

by Newey (1985) for constructing conditional moment tests.[10] Thus we test $H_0 : \gamma_1 = \gamma_2 = 0$ in

$$P(Dep = 1 \mid x) = \Lambda(\beta^T x + \gamma_1(\beta^T x)^2 + \gamma_2(\beta^T x)^3)$$

using the test statistic proposed by Newey (1985), adapted to the logit specification. We obtain as test result for the no-privatisation case a value of 6.16; since this test is asymptotically chi-squared-distributed with 2 degrees of freedom, the test leads to rejection of the logit specification at the 5%-level. In case of the in-house modelling the test result becomes much higher: 33.95; this means rejection of the logit specification at all usual significance levels.

Consequently, it makes sense to investigate alternative specifications, which require less severe distributional assumptions. One possibility is a fully nonparametric approach, but due to the curse of dimensionality this will not work in our case with only 540 observations. So, we restrict attention to semiparametric models. There are several possibilities available in the literature for application to the binary choice case. One possibility is the Maximum Score estimator proposed by Manski (1985), and turned into smoothed Maximum Score by Horowitz (1992). Although (Smoothed) Maximum Score requires very weak distributional assumptions it has some drawbacks: it has a lower rate of convergence than ordinary parametric estimators and it only allows one to estimate the index, but not the probability transformation. Another possibility are the single index models in which the probability that the binary dependent variable equals one given the covariates is equal to a single index of the covariates evaluated in an unrestricted (nonparametric) probability transformation:

$$P(Dep = 1 \mid x) = H(\beta^T x),$$

where H is an unknown function that has to be estimated as well. There are several estimators available to estimate such single index models. For instance, Klein and Spady (1993) provide a semiparametric efficient one. However, this estimator is quite hard to calculate in practice. We decided to use Ichimura (1993).[11] The estimator for β consists of solving the minimisation problem

$$\hat{\beta} = Arg \min_b \sum_i (Dep_i - \hat{H}(b^T x_i))^2,$$

where $\hat{H}$ represents a nonparametric estimator for

$$P(Dep = 1 \mid x) = E(Dep|x) = H(\beta^T x).$$

We estimate this latter conditional expectation using a kernel estimator with a standard normal Gaussian kernel. Since there is no optimality theory for the corresponding bandwidth, we have set it equal to the familiar rule of thumb $\hat{\sigma} n^{-1/5}$, with $\hat{\sigma}$ an estimate for the standard deviation of $\hat{\beta}^T x$.[12] The resulting

[10] Newey (1985) considers the probit specification; however, the adaptation to the logit model is straightforward.

[11] For other possibilities, see, for instance, Horowitz (1998).

[12] As a starting value for the iteration procedure we used the OLS-estimate for β, from which we also constructed the estimate for σ.

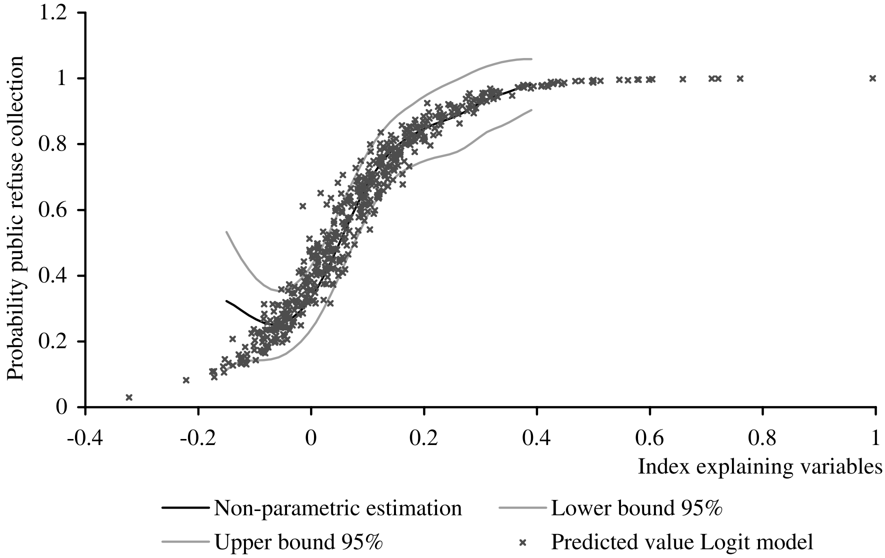

Fig. 1. Non-parametric (solid lines) and logit (dots) estimation of choice between public and private refuse collection

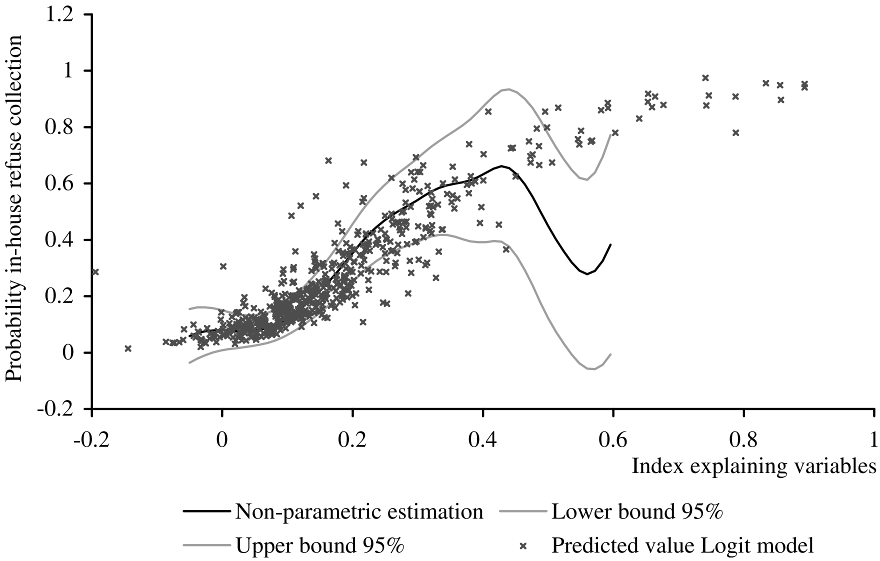

Fig. 2. Non-parametric (solid lines) and logit (dots) estimation of choice between in-house and out-house refuse collection

estimator for β has a normal limiting distribution whose asymptotic covariance matrix can straightforwardly be estimated. See Ichimura (1993) for further details.

Table 2 contains the estimation results for β, and Figures 1 and 2 present the estimates for H, for the no-privatisation and in-house case, respectively. Notice that in the single index model the constant term is not identified (therefore, set equal to 0). Also the scale is not identified; we have fixed the scale by normalising the coefficient of the variable Inhabitants, equal to the corresponding estimated coefficient in the Logit model.

The estimation results in terms of β according to Ichimura are, at least qualitatively, quite comparable with those according to the logit specification. To investigate whether the results are also quantitatively the same, we considered the hypothesis that the coefficients of logit are (simultaneously) equal to the corresponding single-index coefficients of the Ichimura-specification. We tested this hypothesis by a Hausman-type test by using the difference of the vector of logit estimates and the corresponding Ichimura-estimates. The limit distribution of this difference can easily be obtained under the null hypothesis. The value of the resulting chi-square test statistic turned out to be 1.75 in case of no-private collection, and 3.69 in case of in-house provision. Since, under the null hypothesis, the test statistic is asymptotically chi-square-distributed with 13 degrees of freedom (the number of coefficients, except the constant term and the normalised coefficient of inhabitants), we conclude that the results in terms of the single-index coefficients are also quantitatively the same.

Next, we turn to the estimated probability transformations. In Figure 1 we plot the nonparametric estimate of the probability transformation in case of no-private-collection, together with 95% confidence intervals. In addition, we plot in the figure the corresponding predictions according to the logit model. From this figure we can conclude that the logit- and the Ichimura-specifications for most observations are not too far apart from each other. However, a non-negligible part of the predictions according to logit fall outside the 95% confidence interval, which can be seen as evidence that the logit model is misspecified, in line with the earlier rejection of the logit probability transformation. Moreover, for the lowest values of the single-index the results of Ichimura differ substantially from logit, although not significantly so. It seems that the probability transformation is not increasing over the whole range, a feature that cannot be captured by the logit-specification. In Figure 2 we present the corresponding plot in case of in-house-provision. Again, we see that for many observations the logit- and Ichimura-specifications are reasonably close, but not as close as in case of no-private-collection: Over the whole range we see predictions according to logit falling outside the 95%-confidence band.[13] Moreover, for larger values of the single index, the Ichimura probability transformation is not increasing, but inversely hump shaped, a pattern that clearly cannot be captured by the logit probability transformation. Concluding, based on the overall evidence, the difference between logit and Ichimura is significant, in line with the earlier reported rejection of the logit probability transformation.

To investigate the consequences of the misfit by logit for a substantial part of our sample, we compare the prediction performances of the models, as well as the estimated marginal effects of changes in the covariates on the probabilities. First, Table 3 contains the prediction performances. For the sake of comparison, we also include in this table the naïve predictions without using

[13] The number of inhabitants, population density and the share of local parties deviate for the municipalities outside the 95%-confidence band. Probably the Ichimura specification is especially superior for observations with special characteristics as this specification allows more flexibility.

Table 3. Percentage correct predictions of the various models

	No-private collection	In-house collection
Naïve prediction	0.58	0.72
Logit	0.7574	0.7685
Ichimura	0.7593	0.7796

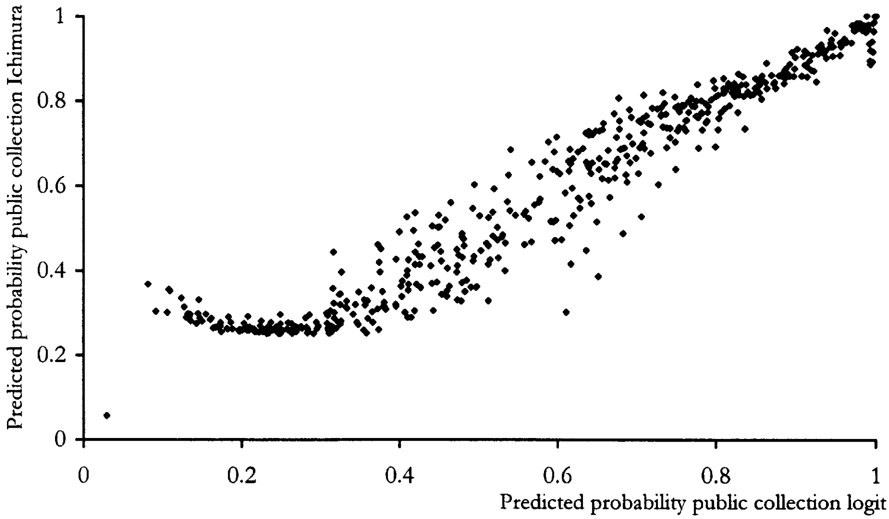

Fig. 3. Comparison predictions logit versus Ichimura (public versus private refuse collection)

any covariates. We predict the endogenous variable to be equal to one, if the predicted probability is at least a half; otherwise, we predict the endogenous variable as zero.

From this table we conclude that the prediction capabilities of both logit and Ichimura are quite comparable, and that Ichimura only slightly outperforms logit in both the no-private-collection and the in-house-collection cases. Of course, this is only a very rough comparison. To further illustrate how the predictions of both specifications are in line with each other we plot in Figures 3 and 4 the predictions according to logit against those according to the Ichimura specification. Figure 3 contains the comparison for the no-private-collection case. The correlation coefficient between the predictions equals 0.97, but, particularly, at lower values of the single indices, we see a clear difference between both specifications, as already suggested by Figure 1, but not reflected in Table 3. In Figure 4 we consider the in-house-collection case. Here the correlation is much weaker than in Figure 3. Indeed, the correlation coefficient is only 0.87, indicating that a blind use of logit may be misleading.

Thus, although the single index of logit corresponds quite closely to the single index according to Ichimura, the logit probability transformation is likely to be misspecified, due to its inflexibility, preventing it from fitting non-monotonic patterns. This might have implications for the quantification of the marginal effects of the covariates on the probabilities of no-private collection and in-house provision. To investigate this, we compare the estimated mar-

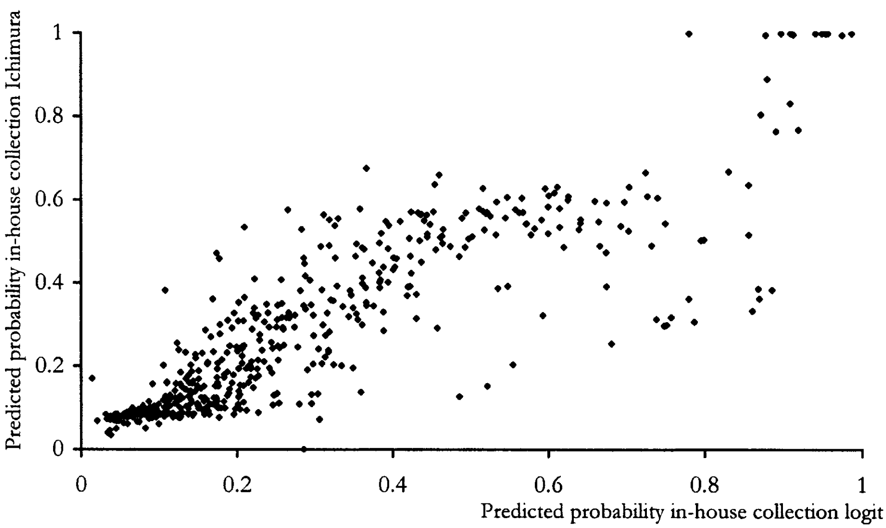

Fig. 4. Comparison predictions logit versus Ichimura (in-house versus out-house)

ginal effects of changes in the covariates on the predicted probabilities. We calculate these effects for each municipality in our sample, and then we average them over the sample. In this way we are measuring the (average) macro-effect of a marginal change in the covariates. Notice that our sample contains almost all Dutch municipalities, so that we are more or less dealing with the whole population. We include the standard deviations of the means to give an indication of the variability of the calculated effects, and we calculate the average of the absolute differences per municipality between the two models, to see how close the effects are. Table 4 contains the results for no-private-collection and Table 5 presents the results for in-house-collection.

Looking first at Table 4 (no-private-collection), we see in case of, for instance, the output variables (inhabitants or inhabitants per hectare) that the calculated average macro effects are quite comparable between the two specifications. However, in both cases, the average absolute differences are quite large compared to the average marginal effects, indicating that on the individual municipality level the models yield substantial differences, which, on an aggregate level, are averaged out. We also see that the variability in the logit marginal effects is much smaller than the variability in the Ichimura marginal effects, which, of course, is a consequence of the imposed monotonic logit probability transformation, as opposed to the non-monotonic Ichimura probability transformation. In case of the wealth variables and the interest group variables we have a similar story. Looking at the political variables, we see that in some cases the differences between both models are, at least qualitatively, substantial, although the magnitudes of the marginal effects are quite small.

Turning next to Table 5 (in-house-provision), we see that the differences are now more substantial. For instance, in case of the output variables the estimated marginal effect according to Ichimura is between 1.6 (inhabitants) and 2.6 (inhabitants per hectare) times as large as the corresponding effect according to logit. In case of the wealth variable income per inhabitant

Table 4. Marginal effects no-private-collection (standard deviations in brackets)

Variable	Logit	Ichimura	Abs. Difference
Inhabitants	0.0437	0.0484	0.0269
	(0.0182)	(0.0425)	
Population density	0.0133	0.0135	0.0072
	(0.0053)	(0.0117)	
Funds	0.2413	0.2165	0.1101
	(0.0967)	(0.1871)	
Income	−0.2815	−0.2420	0.1223
	(0.1127)	(0.2091)	
Unemployment	0.0037	0.0035	0.0018
	(0.0015)	(0.0030)	
Civil servants	0.0471	0.0377	0.0191
	(0.0189)	(0.0325)	
Conservative Liberals	−0.0007	−0.0023	0.0020
	(0.0003)	(0.0020)	
Social Democrats	−0.0001	−0.0035	0.0038
	(0.0000)	(0.0030)	
Progressive Liberals	0.0014	−0.0030	0.0045
	(0.0006)	(0.0026)	
Orthodox Protestants	0.0047	0.0022	0.0026
	(0.0019)	(0.0029)	
Green Left	−0.0033	−0.0082	0.0065
	(0.0013)	(0.0071)	
Extreme Right	−0.0081	−0.0172	0.0131
	(0.0032)	(0.0148)	
Local Parties	−0.0041	−0.0062	0.0040
	(0.0016)	(0.0053)	

the estimated negative marginal effect of income per inhabitant in case of Ichimura is even almost three times as large as in case of logit. The corresponding average absolute differences are also quite substantial. Similarly to the no-private-collection case, the logit marginal effects again show much less variability than the Ichimura marginal effects.

Concluding, we can state that, although the logit single index seems to be appropriate, the logit probability transformation seems to be too inflexible, producing, at least in the in-house provision case, average marginal effects whose magnitudes may be quite incorrect, and resulting in both the no-private collection and the in-house provision cases in an accuracy which may be quite misleading. By applying a semiparametric specification, this inflexibility of the logit probability transformation can easily be circumvented.

So far, we considered no-private-provision and in-house-collection separately. However, one might argue that there may be some ordering present: at level 0 one can consider full privatisation; at level 1 there is public provision, but not in-house; and at level 2 there is full in-house collection. Such an ordering may be modelled by a single index model as well. However, this only makes sense if the two indices, when estimating the choices no-private-collection and in-house-provision separately, are (more or less) the same. Therefore, we also considered the hypothesis that the vectors of coefficients of these two indices are equal. We tested this hypothesis by means of a Hausman-type test based on the difference between the two Ichimura-

Table 5. Marginal effects in-house-collection (standard deviations in brackets)

Variable	Logit	Ichimura	Abs. Difference
Inhabitants	0.0257	0.0426	0.0314
	(0.0110)	(0.0432)	
Population density	0.0051	0.0136	0.0121
	(0.0022)	(0.0149)	
Funds	0.2092	0.3378	0.2596
	(0.0890)	(0.3702)	
Income	−0.1029	−0.2934	0.2659
	(0.0438)	(0.3215)	
Unemployment	0.0258	0.0109	0.0083
	(0.0067)	(0.0120)	
Civil servants	0.0157	0.0122	0.0084
	(0.0067)	(0.0123)	
Conservative Liberals	0.0001	0.0022	0.0025
	(0.0001)	(0.0025)	
Social Democrats	−0.0063	−0.0066	0.0045
	(0.0027)	(0.0072)	
Progressive Liberals	0.0042	0.0029	0.0022
	(0.0018)	(0.0032)	
Orthodox Protestants	0.0037	0.0024	0.0020
	(0.0016)	(0.0026)	
Green Left	−0.0018	−0.0034	0.0028
	(0.0008)	(0.0038)	
Extreme Right	−0.0270	−0.0721	0.0642
	(0.0115)	(0.0790)	
Local Parties	−0.0041	−0.0020	0.0024
	(0.0018)	(0.0022)	

estimators, after appropriate scaling.[14] The resulting chi-square test statistic yielded as value 36.5, which results in strong rejection of the hypothesis of equal indices, since the critical value of a chi-square distribution with 13 degrees of freedom equals 22.36 (at 5%). We concluded that the modelling of the mentioned ordering by means of a single index is likely to yield a misspecified model, even if modelled semiparametrically. Therefore, we did not investigate this possibility further.

6. Conclusions

In this paper we try to explain the reasons why contracting out refuse collection is less common than in-house provision, although considerable efficiency improvements by contracting out seem achievable. We present an empirical investigation motivated by output arguments, interest group theory, and ideology arguments.

We used both a parametric (logit) and a semiparametric (Ichimura (1993)) modelling approach, which correspond in the use of a single index, but which differ in terms of the flexibility of the probability transformations employed. The estimated single indices are quite similar, so that both yield the same

[14] The limit distribution of this difference can easily be obtained under the null hypothesis.

conclusions, when investigating the direction and statistical significance of the various effects.

In both models we find evidence for the hypothesis that a high level of transfers by the central government (the wealth argument) or a high level of unemployment (the interest group argument) raises the probability of public and in-house provision. We also find evidence for the assumed relation between the size of municipalities and private collection. In all cases a smaller municipality is more likely to have private collection. Therefore, scale effects are important for the choice between public and private provision. For the choice between out-house and in-house collection in relation to scale lesser evidence exists. Weak evidence is found for an ideological motivation of this choice.

However, when explicitly quantifying the size of these effects, one also needs the probability transformation, transforming the single index into the probability that the dependent variables equals one. Here, we find strong statistical evidence that the parametric specification is far too inflexible, with the danger that the corresponding estimated marginal effects might be misleading. Indeed, in a number of cases, we find serious differences between the parametric and the semiparametric marginal effects, implying that one should be very cautious, when using parametric models.

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