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# The link between service privatization and price distribution among consumer types: municipal water services in the Spanish region of Catalonia

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**Abstract.** In the Spanish region of Catalonia the industrial sector sensibly lobbied for the privatization of municipality-produced water services in the period 1996–2002. This paper analyzes the extent to which this result is consequent with observed measures of cross-subsidies in favor of residential consumers. I run a treatment-effects regression on a sample of 133 municipalities in 2000 and 2001. Recent privatization has a positive effect on the weight that the local regulator attaches to industrial users' welfare when deciding water tariffs. I conclude that an increase of industrial customers' group size has a twofold effect on the cross-subsidy in favor of residential customers: a direct *positive substitution effect* and an indirect negative *political support effect* through lobbying for service privatization.

## 1 Introduction

According to my earlier work (Miralles, 2006), there are several factors influencing local water-service privatization in Spanish municipalities<sup>(1)</sup> during its current democratic period (1979–2002). In fact, these factors change over time, as I found in my results.<sup>(2)</sup> One of the factors tested during this long period was the presence of powerful groups of interest. I focused on the industrial sector, since there usually exists a different tariff structure between residential consumers and industrial ones. It could be suspected that a local regulator could cross-subsidize residential consumers by rising industrial-use tariffs above the socially efficient level. Therefore, in municipalities where the industrial consumers constitute an important part of the water market, they could try to join forces and provoke a change of regime that gives, as a result, a better deal for the industrial customers.

These results indicate that the industrial sector exerted a significant pressure on the water-service privatization decision only during the 1990s, and that this pressure was concreted during the 1996–2002 period. The forecasted advent of the Water Framework Directive (2000/60/European Commission, 23 October 2000)<sup>(3)</sup> and its increasing effects on future water prices may have alarmed industrial consumers, who may have received substantial incentives to join a strong pressure group towards a change of regime.

<sup>(1)</sup> More precisely, my sample covers 133 municipalities in the Spanish region of Catalonia. Results of this paper are then extensive to Catalan municipalities. Other regions in Spain have sensibly different privatization patterns (see Bel, 2006) concerning the local water service.

<sup>(2)</sup> Apart from my own work (Miralles, 2006) the only empirical work I am acquainted with that studies this dynamicity is Hefetz and Warner (2004), for American municipalities. They are particularly interested in the reverse-privatization phenomenon that is growing in the United States.

<sup>(3)</sup> See Buller (1996) for previous European regulation and its effects on water service regulation in several European countries, above all, France and England.

It is not clear, though, why an interest group should push for a policy that indirectly affects the tariff structure instead of directly putting pressure on the tariff formation. This motivates the present analysis, which tries to give robustness to the aforementioned results. If previous results are correctly interpreted, the privatization decision of a local water service in, say, the period 1996–99 must, on average, be linked to an alteration of the tariff policy in favor of industrial customers in further years. This alteration should be more intense than that which is generally observed in the full set of municipalities. As I show in this paper, this appears to be the case.

There is a political economy rationale underlying the result. When the local regulator sets water prices, he or she is maximizing a weighted utility function. The lobbies' pressure has an influence on these weights, as suggested by the general theory of regulation of Stigler (1971) and Peltzman (1976). But altering the status quo between industrial consumers and residential consumers has political costs. Households (that is to say, voters) may not understand why their tariff rise is relatively more than industrial-use tariff rises. Therefore, the local regulator thinks twice before making major changes in the cross-subsidization scheme. In order to minimize the associated political costs, the regulator needs a new environment that could justify major changes.<sup>(4)</sup> This is due to the temporary preeminence of a discourse in favor of service reforms and to the uncertainty surrounding the process. Local water services that have not recently undertaken major regulatory reforms do not have this environment. And privatization is indeed one of the major reforms in the local water service.<sup>(5)</sup>

The reader might think that the finding of empirical evidence in favor of my hypothesis might be related to other explanations besides industrial lobbying. For instance, the municipality could have devised water-service privatization with the aim of pursuing financial balance of the service. According to the hypothesis that cross-subsidies in favor of residential users existed, the consequent rise in water tariffs would typically be higher for residential consumers than for industrial ones. For the reason described in the previous paragraph, such a change would need a context of major service reforms in order to avoid political suicide. Admittedly, this explanation cannot be empirically disentangled from my first hypothesis with the data at hand. However, it is apparent from Miralles (2006) that the industrial lobbies were influential in the water-service privatization decision during the period under study. This gives reliability to the industrial-lobbying approach. In any case, privatization is used to reduce cross-subsidizations, but the pressure to reduce these cross-subsidies could come from different agents.<sup>(6)</sup>

A second point the reader could be suspicious of is whether there is a reason why privatization should be linked to lower political costs of reducing cross-subsidizations.<sup>(7)</sup> Indeed, let me be clear, perhaps redundantly, about this. I neither argue for nor contest the fact that cross-subsidies are lower in municipalities where the service is

<sup>(4)</sup> Giuletti and Otero (2002) analyze the timing of major tariff structure changes in regulated industries in England and Wales. Concerning the water industry, it is observed in the majority of regional water authorities that major tariff structure changes started directly after the 1989 privatization. Ordoñez de Haro (2002) considers the case of several local water utilities in the Spanish southern province of Malaga. Concerning privatization and tariffs, the author finds that observed privatization processes are followed by some "unjustified" price increase.

<sup>(5)</sup> Admittedly, it is difficult to distinguish between these two issues: (1) privatization is a maneuver specifically designed to lead to a costless change in the cross-subsidization scheme, and (2) privatization as a policy that was undertaken for other reasons, but was then taken as a politically costless way to concede to industrial lobby's pressures.

<sup>(6)</sup> I am grateful to a referee for bringing this idea to my attention.

<sup>(7)</sup> Once again, I thank a referee for warning me about this possibility.

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produced by a private supplier than where it is publicly produced. I claim instead that cross-subsidies in favor of residential users are (perhaps temporarily?) reduced in municipalities under a major reform process, privatization being one of them. In my case study the test would be to assess whether recent privatization induces lower price differentials. So I compare those municipalities that have not experienced major reforms for many years to the ones that have done so recently. My hypotheses are: (1) in the long run, equilibrium cross-subsidies do not depend on the regime (private/public production), but (2) shocks on the regime temporarily affect the water tariff scheme.

There is some literature dealing with utility price regulation and subsidies. Usually, it is mainly concerned with measuring the welfare gains that would arise in the presence of a socially efficient regulation (Garcia and Reynaud, 2003; Garcia-Valiñas, 2005; Renzetti, 1999; Timmins, 2002a). Some work has also been done to study cross-subsidization among types of consumers by income. Beard and Thompson (1996) develop a model of a bipartisan two-part tariff voting game. Under some mild conditions they find that a vote-maximizing party will charge a marginal price that is below that which is socially optimal and choose a fixed fee that is above the socially optimal fee. This tariff scheme is usually more favorable to high-income customers. Timmins (2002b) develops and empirically tests a simple model of a bipartisan price–tax voting game. There are two variables to choose: the water marginal price and the local income (wealth) tax rate. He finds that, when municipalities regulate water tariffs, those with a larger income dispersion would charge a (subsidized) marginal price that goes further away from marginal costs, in order to redistribute resources in favor of poorer citizens.

Besides this work there is some work on cross-subsidization among user types. Bruggink (1982) measured the extent of third-degree price discrimination in the municipally owned water industry between industrial, commercial, and residential customers. They find existence of a cross-subsidization *in favor of* industrial customers for a US case in 1965. Ross (1984) develops an easy model to extract regulator's preference weights among consumer types from the so-called Ramsey numbers (ie markups weighted by demand elasticities). Klein and Sweeney (1999) empirically estimate factors affecting these relative weights that the regulator give to industrial, commercial, and residential customers in the gas distribution utilities. They find that the industrial sector size, measured both as the log of industrial consumption and as the log of the number of industrial users, positively affects the relative weight allocated to this sector. The present paper tries to extend some of these ideas to the water-utilities regulation. Nelson and Roberts (1989) apply the same procedure to the electricity-supply market, also finding size effects on the weight attached to the industrial customers. Naughton (1989) develops a more complete model to study regulators' preferences in type-specific two-part tariffs.

In this paper I use the same sample of 133 Catalan municipalities as in Miralles (2006). These municipalities responded to the Local Service Production Survey (LSPS) elaborated by the 'Public Policy and Economic Regulation' research unit at the University of Barcelona. They provided information on the mode of production of the water services and, when applicable, on the year that they were privatized.

Section 2 comments on the regulation of water prices in the Spanish region of Catalonia and describes a simple theoretical model that further justifies econometric estimations of some measure of cross-subsidization between industrial consumers and domestic consumers. Section 3 contains the estimation of the cross-subsidization equation and also an interpretation of the results. Section 4 concludes the paper.

## 2 Water-price regulation: a simple model

In the Spanish region of Catalonia, municipalities propose water tariffs concerning water treatment, water provision, and sewage, whether the service is privatized or not. Municipalities can also associate, forming a regulatory agency if they wish. In the case of municipalities where the service is privatized, either tariff update formulas are included in the contract or prices are renegotiated at a frequency stipulated in the contract. The Catalonia Price Commission receives these proposals and either authorizes them or amends (always downwards, by law) them. The study of this part of the water tariff is the main interest of the present paper, since a vast majority of my sample of municipalities distinguish between industrial users and residential users when fixing the tariff structure.<sup>(8)</sup>

Additionally, the Catalonia Water Agency, created in 1999, has annually fixed the water tax (canon del agua) since April 2000. The water tax distinguishes between nonresidential customers and residential users. Revenue fund works in dam maintenance, environmental control, and water waste treatment plant runnings, among other expenses. Finally, a 7% retail tax is charged over the full amount of the water bill, and the related revenues are shared between the state and the Spanish regions. While these two components of the water bill are not of primary interest, they are important in that they also determine water demand.

To get started with the model, we assume that there are only two types of consumers: industrial (I) and residential (R). For each one a semilog demand function with respect to marginal price is modeled. I refer the reader to Timmins (2002a) who justifies this functional form using several arguments. First, it allows for a nonzero demand for any price, which makes sense because water is a necessity good, even for industrial customers.<sup>(9)</sup> Second, it has a satiation point: demand does not go towards infinity when the price goes to zero. Again, this makes sense for both types of consumer. Both the marginal utility and the marginal productivity of water use could become negative if too much water is consumed. In the present model this functional specification is also good because the semilog specification implies constant price semielasticity, and this simplifies further calculations. Therefore, the respective (individual) demand equation is:

$$\ln \text{DEM}_{jit} = \text{NPD}_{jit} - \text{SE}_j \text{TMP}_{jit}, \quad j \in \{I, R\}, \quad (1)$$

where  $i$  refers to municipality,  $t$  to year, and  $j$  to consumer type.  $\text{SE}_j$  is the semielasticity for consumer type  $j$  with respect to TMP, which is the total marginal price ( $\text{TMP} = \text{MP} + \text{TAX}$ , where MP is the municipality-proposed water marginal price and TAX is the marginal rate of the water tax) ignoring retail taxes.<sup>(10)</sup> NPD refers to the component of demand that is not price dependent. Let us denote  $\text{SUR}_{jit}$  for individual consumer-type  $j$ 's surplus in municipality  $i$  at time  $t$ , depending on  $\text{TMP}_{jit}$ . Then, the local regulator, when setting his or her water tariff structure, solves the following equation:

$$\begin{aligned} & \underset{\text{MP}_{Rit}, \text{MP}_{Iit}}{\text{maximize}} \text{WRC}_{it} \text{RAT}_{it} \text{SUR}_{Rit} + (1 - \text{WRC}_{it} - \text{WSP})(1 - \text{RAT}_{it}) \text{SUR}_{Iit} \\ & + \text{WSP}[\text{RAT}_{it}(\text{MP}_{Rit} - \widetilde{\text{MC}}_{Rit})\widetilde{\text{DEM}}_{Rit} + (1 - \text{RAT}_{it})(\text{MP}_{Iit} - \widetilde{\text{MC}}_{Iit})\widetilde{\text{DEM}}_{Iit}], \quad (2) \end{aligned}$$

<sup>(8)</sup>In some cases there is a specific tariff for commercial users. This sector is not analyzed here due to the lack of related data. Also, there is almost always a municipality-use subsidized tariff, which is not of interest to us here.

<sup>(9)</sup>Martinez-Espineira and Nauges (2004) develop a Stone–Geary demand function for water, where there is a minimum consumption level that is insensitive to prices. This is also a good way to model for necessity goods.

<sup>(10)</sup>The semielasticity of interest for the municipal regulation must include the retail tax multiplicative effect, because the one-currency-unit municipal marginal-price increase is amplified by the retail tax.

subject to

$$MP_{Rit}, MP_{Iit} \geq 0 .$$

The  $\sim$  symbol can be taken to mean ‘belief or signal about the true value’. MC is the [constant<sup>(11)</sup>] marginal cost; WSP is the weight allocated to the service producer and is assumed to be constant for the sake of simplicity. There is some intuition in this assumption. The Catalonia Price Commission can amend the proposed local water tariff downwards, although it seldom modifies the cross-subsidy scheme between types of consumers. Therefore, MC can be understood to be a parameter set finally by the Catalonia Price Commission, and hence to be common to all municipalities in the studied sample.  $WRC_{it}$  is the weight that the municipal regulation attaches to residential customers and  $RAT_{it}$  is the ratio of the number of residential users over the sum of residential and industrial users.

Notice that in this setup it is assumed that service fees are chosen previously in order to cover for fixed costs, as in Timmins (2002a). They could also be set as the result of any other separate process. The reader might wonder whether this is a good procedure. Naughton (1989) develops a model in which marginal prices and fixed fees are jointly set as the solution to a more complete maximization problem. This way he is able to identify not only intertype relative weights but also intratype weights (eg, between poor residential users and rich residential users). There are several reasons, though, why I do not use his approach. First, concerning my sample, not only are fixed fees type variant but also there are differences inside the same consumer type. The resulting fixed-fees structure differs from municipality to municipality in a quite messy way. There is no good measure for what is understood to be an average type-specific fixed fee. Second, Naughton says that the model presented here is equivalent to his model if two conditions hold for each consumer type: (1) the number of potential users already connected to the utility is not sensitive to the tariff structure, and (2) the demand elasticity with respect to the fixed fee is zero. In my sample, quoting data from Idescat, more than 95% of residential buildings in Catalonia are connected to a municipal water utility, as of 2001.<sup>(12)</sup> Only low-populated villages have higher percentages of self-supplied households, and this is a phenomenon that cannot be well explained by abusive water tariffs. And I will also assume that residential users do not ‘vote with their feet’ concerning water tariffs, so that the household-location decision is not a function of the water tariffs. Therefore, the number of residential users in a municipality is rather insensitive to the water tariff. A last observation is that the fixed fee is such a tiny part of residential user’s income that the demand elasticity is negligible with respect to the fixed fee. Concerning the industrial users, I will assume that there are factors playing a much bigger role in the location decision than the water tariff.<sup>(13)</sup> Moreover, it is generally accepted that the industrial customers’ water demand is not sensitive to fixed fees.

First-order conditions yield the following equations:

$$MP_{Rit} = \max\{0, LP_{Rit}\}, \quad LP_{Rit} = \widetilde{MC}_{Rit} + \widetilde{SE}_R^{-1} \left( 1 - \frac{WRC_{it}}{WSP} \right), \quad (3)$$

<sup>(11)</sup> Assuming constant marginal costs with respect to quantity is not at odds with what has been observed in the literature. Trujillo (1994) supports this hypothesis with real data.

<sup>(12)</sup> Similar statistics stand for the sewerage service.

<sup>(13)</sup> There are specific industrial sectors that give importance to the water tariff, since they are big water consumers. The representative firm’s strategy in these sectors could consist of locating where water prices are low, sacrificing other location features. Or it could consist of locating where other features are appealing, lobbying for water-price reduction thereafter. In this paper it is shown that industrial lobbying is quite successful under some conditions, so the location decision may not be very sensitive to original water tariffs, even in these sectors.

and

$$MP_{Iit} = \max\{0, LP_{Iit}\}, \quad LP_{Iit} = \widetilde{MC}_{Iit} + \widetilde{SE}_I^{-1} \left[ 1 - \frac{(1 - WRC_{it} - WSP)}{WSP} \right]. \quad (4)$$

LP refers to latent prices that would be set if the nonnegativity constraints had not existed. If latent prices were nonnegative and the local regulator's beliefs about marginal costs were known by the econometrician, the relative weight  $WRC_{it}/(1 - WRC_{it} - WSP)$  could easily be calculated and regressed on a series of political, social, and other variables, as suggested in Ross (1984)<sup>(14)</sup> and analogous to Klein and Sweeney (1999) and Nelson and Roberts (1989). Unfortunately, this is not the case in my sample. There are no broadly available data on local water-service costs in Catalonia. And even if there were available data, these might not be reliable. Since an important part of the municipalities of the sample that I analyze has privatized the water service, the private firms may have incentives to misreport cost data in order to obtain better tariff renegotiations. So the beliefs that the regulator has about marginal costs would be unobservable in some cases even if data on accounting costs were widely available. Let us now assume that prices are equivalent to latent prices. We can manipulate the first-order conditions in a different way to obtain

$$\Delta MP_{it} = \Delta \widetilde{MC}_{it} + \Delta \widetilde{SE}^{-1} - \widetilde{SE}_I^{-1} \frac{(1 - WSP)}{WSP} + \frac{WRC_{it}}{WSP} \sum \widetilde{SE}^{-1}, \quad (5)$$

where  $\Delta MP_{it} \equiv MP_{Iit} - MP_{Rit}$ ,  $\Delta \widetilde{MC}_{it} \equiv \widetilde{MC}_{Iit} - \widetilde{MC}_{Rit}$ ,  $\Delta \widetilde{SE}^{-1} \equiv \widetilde{SE}_I^{-1} - \widetilde{SE}_R^{-1}$  (presumably negative, since residential demand for water tends to be more inelastic than industrial demand), and  $\sum \widetilde{SE}^{-1} \equiv \widetilde{SE}_I^{-1} + \widetilde{SE}_R^{-1}$ . This equation enlightens us about several points. The difference of marginal prices between consumer types does not only depend on the difference between respective regulator's beliefs about marginal costs; it also depends on other factors: the difference in inverse semielasticities, the relative preference for consumers against the service producer, and the relative preference in favor of one type of consumer.

The price differential equation [equation (5)] is the base for my estimations. The  $WRC_{it}/WSP$  factor is approximated by a function of different political factors, among them the recent privatization decision.

### 3 Main estimations

Equation (5), is linear in  $WRC_{it}/WSP$ , which is the regulator's revealed weight in favor of residential users (with the assumption that the latent prices are positive). This equation has pros and cons. On the negative side, the assumption that observed marginal prices coincide with latent prices is too strong in my sample, since there is a substantial amount of zero-price observations. Due to this reason a bivariate Tobit maximum-likelihood model has been constructed and estimated. The Tobit-like model is robust to a relaxation of the previous assumption. Results are not presented here, although they are available upon request. Concerning the variable of interest in this paper (that is, recent privatization), conclusions are unchanged, so the bias implied by this assumption can be considered to be negligible. The second drawback of the proposed equation is that, in order to obtain a parsimonious model, I have

<sup>(14)</sup> Ross warns about the use of this formula when the econometrician is dealing with block-rate pricing, which is the case in the sample. Although I admit this, no better theoretical approach is easily implementable. I will just assume that the subsidies implied by increasing block rates (the so-called difference variable) are chosen along with the fixed fees.

assumed that WSP, the service producer's weight in the regulator's objective function, is constant.<sup>(15)</sup> We will see that this parameter is not easily identifiable.

On the positive side it is worth mentioning that we do not need to approximate the regulator's beliefs about the common components of the different marginal costs. Given the lack of cost data, there would be almost unavoidable approximation errors—consider the different sources of water, water scarcity, the price paid to a possible supramunicipal water provider, wages, interest rates, electricity prices, management costs etc. And, furthermore, it is not only the marginal costs but also the regulator's beliefs in them that must be considered. This is a very important consideration when the service has been privatized through a concession contract—the most usual way in Catalonia. When tariff renegotiation is coming the firm reveals misleading cost information if this is going to be profitable to itself, and the regulator has to guess to which extent the information given is biased. The previously mentioned double Tobit model has to deal with this problem. That is why a more complete model may not perform better in this case, unless reliable data on costs are available. The simplicity of the proposed model can substantially offset the bias problem.

So, I have a simpler task: I need to approximate the regulator's beliefs about the difference in marginal costs. Therefore, I simply need to approximate the components of one marginal cost that is not part of the other marginal cost in the regulator's eyes. This implies that I must forget about common components of marginal costs such as those mentioned above.

I model the difference between regulator's beliefs about differences in marginal costs in a very parsimonious way:

$$\widetilde{\Delta MC}_{it} = \beta_0 + \beta_1 B3_i + \beta_2 IRB_i + \beta_3 TT + \text{ERROR}_{it}^A . \quad (6)$$

B3 is the proportion of residential buildings in the municipality with more than three floors, as of the year 2000. Data have been collected from Idescat. The height of residential buildings is supposed to (positively) affect residential water marginal cost only. Therefore, the associated coefficient is supposed to be negative. I also include a geographic dummy variable, IRB. It accounts for the internal river basins of Catalonia—rivers whose full length is contained within this region. This accounts for unobservables related to geographical location.  $\text{ERROR}_{it}$  is an error term. Finally, TT is a time trend, taking a value of 1 if the year is 2001 and 0 if the year is 2000.

I additionally model residential users' relative weight  $\text{WRC}_{it}/\text{WSP}$  in a rather parsimonious way, as a function of a set of social and political variables:

$$\frac{\text{WRC}_{it}}{\text{WSP}} = \gamma_0 + \gamma_1 M9903_i + \gamma_2 \text{CON}_i + \gamma_3 \text{ILS}_{it} + \gamma_4 P9699_i + \text{ERROR}_{it}^B . \quad (7)$$

M9903 is a dummy that takes the value of 1 if the mayor during the 1999–2003 term was left oriented, and 0 if the mayor was right oriented. The source for this data consisted of the municipality files compiled by the Catalonia Municipalities Federation (FMC). There are many important political parties in Catalonia and also many independent candidates in the municipal elections. It is not always the case that the ideological side of a governing coalition or group is identifiable, above all for small municipalities, so some observations have been lost. The expected coefficient on this variable is positive. CON is a constituency variable, measuring median voter's social preferences. It is constructed as a dummy that takes the value of 1 if the number of left-oriented votes in the two national elections, undertaken in 1996 and 2000, respectively,

<sup>(15)</sup> Although this has been previously justified in section 2.

represent more than half of total votes (data from Idescat). The associated coefficient is assumed to be positive.

ILS is an industrial lobby size variable, measured as the ratio between the number of industrial-water accounts and the sum of residential plus industrial accounts. The data for this variable have been provided by the Catalonia Water Agency. Following Klein and Sweeney (1999) the size of the industrial sector negatively affects the weight that the regulator puts in for other types of consumers. This is the evidence that they find in their sample, although they admit that an increase in the industrial-sector size has several offsetting effects. There is a *substitution effect*, where an increase in the industrial water tariff yields higher revenues, which in turn can yield higher cross-subsidies. And there is a *political support effect*, in which the regulator has to put more weight on the industrial sector's welfare because their support has become more important. This political effect can be made to disappear by an excess of interest-group size, due to coordination failures.

P9699<sub>*i*</sub> is the variable that I am interested in.<sup>(16)</sup> It is a dummy variable that takes a value of 1 if the municipality undertook a privatization process on the water services during the 1996–99 period.<sup>(17)</sup> Apparently, this variable should not have any effect on the revealed preference variable. But in this case, when the European Water Framework Directive is to be implemented, and when the industrial lobbies may put much pressure on the regulator in order to obtain better tariffs, the recent privatization process could be the perfect excuse for the regulator to concede to industry's demands while avoiding the political cost of altering the status quo between residential consumers and industrial consumers. The confusion around how the new tariff structure should be, after the privatization process, and the temporarily dominant discourse in favor of regulatory reforms, may allow the regulator to alter this status quo without cost. So the coefficient attached to this variable can be expected to be negative.

We can estimate the following equation:<sup>(18)</sup>

$$\Delta MP_{it} = \left[ \beta_0 + \theta_0 + \Delta \widetilde{SE}^{-1} - \widetilde{SE}_1^{-1} \left( \frac{1 - WSP}{WSP} \right) \right] + \beta_1 B3_i + \beta_2 IRB_i + \beta_3 TT + \theta_1 M9903_i + \theta_2 CON_i + \theta_3 ILS_{it} + \theta_4 P9699_i + ERROR_{it} . \quad (8)$$

Here,  $\theta_k = \gamma_k \sum \widetilde{SE}^{-1}$ ,  $k = 0, \dots, 4$ . I should point out that there are municipalities in the sample that jointly set the same water tariffs. Therefore, only one observation is permitted per municipality regulatory association.<sup>(19)</sup> From the structure of this equation it can be seen that neither the WSP parameter nor the constants can be identified, although for my purposes this is not an important issue.

I have data on water marginal prices for this sample on the years 2000 and 2001.<sup>(20)</sup> I have elaborated on these data from the edicts that are published by the Catalonia Price Commission when it approves or amends water tariffs. I have taken an annual average marginal price for the average consumer of each type in each municipality

<sup>(16)</sup> Data collected from the LSPS.

<sup>(17)</sup> It is reminded that I found (Miralles, 2006) that the industrial sector only exerts positive influence on the privatization decision from 1996 onwards, whereas data on prices concern the years 2000 and 2001.

<sup>(18)</sup> In this section prices are both deflated using the same index, namely the consumer price index, in order to be comparable.

<sup>(19)</sup> The main municipality of the association in population terms is the municipality of reference when taking explanatory variables data. Exceptionally, the sociopolitical variable CON is based on a weighted average of the whole association of municipalities.

<sup>(20)</sup> The unit used is 2000 pesetas per cubic meter.

and year.<sup>(21)</sup> In doing so I have assumed that water consumption is spread smoothly throughout the year, and that there is no huge variability in consumption among the different consumers of the same type.

There is an important concern here: some endogeneity problems could exist regarding the variable of interest, that is, the recent privatization dummy. The current difference in prices may depend on unobservables that are persistent over time, so that past values of the price difference, which are unobservable here, are included in the error term. But the recent privatization decision can be characterized by the following function:

$$P9699_i = \begin{cases} 1, & \text{if } Z9699_i'COEF + ERR_i \geq 0, \text{ and } PBE96_i = 0, \\ 0, & \text{otherwise.} \end{cases} \quad (9)$$

$PBE96_i$  is a dummy that takes a value of 1 if the municipality privatized the water services before 1996.<sup>(22)</sup> If that is the case there is no further privatization decision and the  $P9699_i$  variable takes the value 0.  $Z9699_i$  is a vector of observed variables influencing the privatization decision during the 1996–99 period.  $COEF$  is a vector of associated coefficients, and  $ERR_i$  is an error term. Since we do not have data on previous values of the price difference, it is quite likely that these are somehow imbedded in this error term. A high cross-subsidization in favor of residential users gives the industrial customers a high incentive to engage in lobbying activities enhancing privatization, as a way to gain some control over the tariff structure. This effect is greatly amplified by the proximity of a new regulation like the Water Framework Directive, which is believed to be going to induce higher tariffs for all customers. Therefore, I could presume that there exists some positive correlation (namely  $\rho$ ) between  $ERROR_{i,t}$  and  $ERR_i$ . Consequently, there probably is some positive correlation between  $P9699_i$  and  $ERROR_{i,t}$ , leading to an overestimation of the associated coefficient  $\theta_4$ .

In order to face this endogeneity risk I have proceeded to a three-part estimation. In the first estimation I simply regress equation (8) by OLS, correcting the standard errors by the use of the White procedure. This way, estimated standard errors are robust to heteroskedasticity. In the second estimation I substitute  $P9699_i$  by an instrument  $IP9699_i$ , namely the estimated probability of privatization during the 1996–99 period calculated from the results in Miralles (2006). In that study I assume that  $-ERR_i$  follows a complementary log–log distribution,  $F(y) = 1 - \exp[-\exp(y)]$ . Robust standard errors are calculated again.

In the third estimation I use a treatment effects model that combines equation (8) and the equation describing the recent privatization decision [equation (9)]. In this model  $Z9699_i$  is substituted for a vector of variables that exerted a significant influence on the privatization decision during the 1996–2002 period, following Miralles (2006). These are: population (negative effect), industrial-sector strength (positive), left-oriented mayor (negative), and the static-neighboring effect (positive). The reasoning underlying the use of each variable has been detailed in Miralles (2006). The first variable,  $POP_i$ , is taken for the year 2001 from the Idescat database. The second variable,  $ISS_i$ , the ratio of industrial customers over the sum of industrial plus residential customers, is taken for the year 2000 by the use of data provided by the Catalonia Water Agency. The third variable  $M9599_i$ , is analogous to  $M9903_i$  but applied to the 1995–99 term of office. The last variable,  $NEI_i$ , is calculated in the following way: for each municipality, the official territory<sup>(23)</sup> that it belongs to is identified; for each territory, the percentage of

<sup>(21)</sup> Data on the average yearly consumption levels of each consumer type have been provided by the Catalonia Water Agency.

<sup>(22)</sup> Data collected from the LSPS.

<sup>(23)</sup> Catalonia is currently divided into seven official territories.

**Table 1.** Differences in water-marginal prices between consumer types in Catalonia, 2000–01.

| Variable                           | OLS                   | OLS instrument          | Treatment effects       |
|------------------------------------|-----------------------|-------------------------|-------------------------|
| Constant <sup>a</sup>              | 7.1994<br>(0.80)      | 22.8886<br>(2.29)***    | 20.7005<br>(2.26)***    |
| B3                                 | -0.4620<br>(-1.45)    | -0.7720<br>(-1.87)*     | -0.6538<br>(-1.47)      |
| IRB                                | 37.8263<br>(4.61)**** | 40.4260<br>(4.96)****   | 29.5588<br>(3.91)****   |
| TT                                 | -10.2786<br>(-1.26)   | -10.2438<br>(-1.30)     | -6.7323<br>(-0.99)      |
| M9903                              | -1.0581<br>(-0.12)    | 0.3248<br>(0.04)        | -3.7667<br>(-0.53)      |
| CON                                | 19.3381<br>(1.84)*    | 14.8620<br>(1.47)       | 21.7791<br>(2.27)***    |
| ILS                                | 87.7058<br>(1.56)     | 76.0314<br>(1.33)       | 125.1903<br>(1.93)*     |
| P9699/IP9699                       | 0.5572<br>(0.04)      | -66.1640<br>(-2.81)**** | -76.3720<br>(-6.42)**** |
| POP                                |                       |                         | -0.00003<br>(-1.94)*    |
| ISS                                |                       |                         | 2.9502<br>(1.94)*       |
| M9599                              |                       |                         | -0.4797<br>(-2.38)***   |
| NEI                                |                       |                         | 0.0170<br>(2.17)**      |
| Constant<br>privatization decision |                       |                         | -0.9067<br>(-2.52)***   |
| $\rho$                             |                       |                         | 0.9491                  |
| Wald test $\rho = 0$               |                       |                         | 32.05 [0.0000]          |
| Observations                       | 178                   | 178                     | 172                     |
| $R^2$ /log likelihood              | 0.1166                | 0.1743                  | -958.1865               |
| $F$ test/Wald test                 | 4.51 [0.0001]         | 5.27 [0.0000]           | 60.09 [0.0000]          |

Note. The  $Z$  statistics are in parentheses; the  $P$  values are in square brackets.

<sup>a</sup>Equal to the term in square brackets in equation (8).

\* Significant at the 10% level; \*\* significant at the 5% level; \*\*\* significant at the 2.5% level; \*\*\*\* significant at the 1% level.

sample municipalities that had already privatized the water services before 1999 is calculated.<sup>(24)</sup> The value of  $NEI_i$  corresponds to the official territory that municipality  $i$  belongs to. Finally, the treatment-effect model is estimated by maximum likelihood, with robust standard errors. This time, I will assume that  $-ERR_i$  follows a standard normal distribution, to test for the robustness of my results to different specifications of the error term.

In estimating these equations some observations are lost because not all municipalities in the sample had industrial water consumers. Results for all estimations are reported in table 1.

The correlation between error terms in the treatment-effect model is higher than 0.94 and is highly significant. This result is, in fact, a positive test for endogeneity of the recent privatization variable. Additionally, I have computed the Durbin–Wu–Hausman test of endogeneity of P9699, taking IP9699 as instrument. The test statistic, distributed as a standard normal under the null hypothesis of no endogeneity, takes the

<sup>(24)</sup>Data calculated from the LSPS.

value of 4.06, which is quite above the 1% significance level. If I instead use PBE99, POP, ISS, M9599, and NEI as instruments, the Durbin – Wu – Hausman statistic becomes 4.76. Therefore, the coefficient on the recent privatization variable is most probably biased upwards, although the fact that it is clearly insignificant may seem appealing and intuitive. When the recent privatization is properly instrumented, the associated coefficient is always significantly negative. This is robust, then, to different specifications of the privatization-decision error term. So I cannot reject my initial hypothesis that there is a link between recent privatization decision and a regulator's revealed preference change in favor of industrial users.<sup>(25)</sup>

At this point, it is worth commenting on two issues. The first one concerns the plausibility of the constant marginal costs assumption. I have tested for that. In both OLS and OLS-instrument regressions in table 1 I have included total residential water consumption and total industrial water use.<sup>(26)</sup> If marginal cost in quantity variant, the coefficient related to quantity should be positive. Since these variables are probably endogenous I have used a touristic variable, TOUR,<sup>(27)</sup> and current population as instruments. For both consumption variables the instruments perform very well (the  $R^2$  coefficient is above 0.90 in all cases). After performing two-stage least squares with robust standard errors I have tested for the joint significance of the two coefficients of interest. This test statistic is distributed as a chi-square with two degrees of freedom with a null hypothesis of no joint significance. For the OLS column the test statistic is 0.15 ( $p$  value is 0.93), and for the OLS-instrument column the test yields 0.50 ( $p$  value is 0.78). In both cases there is no evidence contradicting the assumption of constant marginal costs.

A second issue is related to the tariff structure. Although the underlying theoretical model imposes that marginal price does not depend on quantity when marginal costs are constant, the common tariff structure in the sample involves increasing block tariffs. If the derived *difference variable* (the difference between what one pays and what one would have paid if marginal price were constant at the relevant level) is interpreted as a subsidy that is set along with the service fee in a differentiated maximization problem, as it is assumed here in line with Timmins (2002a), then the average consumption for each type of consumer should not have significant impact on the difference in prices analyzed here. To test for the validity of this assumption I have included the log of average household water demand and (the log of) average industrial user's demand as explanatory variables, in both the OLS and the OLS-instrument columns in table 1. Since the variables are clearly endogenous I have instrumented them by the use of TOUR, an indicator of industrial activity (ACT),<sup>(28)</sup> per capita after-tax income,<sup>(29)</sup> annual rainfall,<sup>(30)</sup> and annual average maximum daily temperatures.<sup>(31)</sup>

<sup>(25)</sup> There is an implicit identification assumption here. It is assumed that if privatization improves efficiency, it reduces (average regulator's beliefs on) marginal costs by the same amount for both types of consumption. While the necessity of this assumption is acknowledged, in my opinion it makes sense. The most important efficiency gain from privatization comes from a general improvement that the firm achieves in the global operation of water services.

<sup>(26)</sup> Data provided by the Catalonia Water Agency. The unit used is the cubic meter.

<sup>(27)</sup> It is measured as the capacity of all hotels, hostels, and campsites in the municipality, divided by its total population, and expressed in per thousand units. The data source is Idescat, and data refer to the year 2000.

<sup>(28)</sup> ACT is defined as the portion of Catalan industrial production already undertaken in the municipality, in per-thousand units. Data come from La Caixa Foundation's economic yearbooks.

<sup>(29)</sup> Again, data come from La Caixa Foundation's economic yearbooks.

<sup>(30)</sup> Data provided by the Catalonia Water Agency, and measured in units of mm/m<sup>2</sup>. Data are only available at *comarca* (county) level.

<sup>(31)</sup> The measure unit is °C. Once again, the data source is the Catalonia Water Agency and data are available at *comarca* level.

Instruments perform well in the case of residential consumption (the partial  $R^2$  is approximately 0.09 in both cases) and quite well in the case of industrial consumption (the partial  $R^2$  is around 0.38 in both estimations). I have performed two-stage least-squares regression with robust standard errors. For the OLS column the Hansen test of overidentifying restrictions yields 1.54 ( $p$  value is 0.67), and for the OLS-instrument case the test yields 1.71 ( $p$  value is 0.63). There is no evidence against exogeneity of the instruments. Finally, the crucial tests, distributed as chi-square with two degrees of freedom under the null hypothesis, yield the following: 0.28 ( $p$  value of 0.87) for the OLS column and 0.78 ( $p$  value of 0.69) for the OLS-instrument column. There is no evidence against the null hypothesis that (log of) per-user consumption levels do not affect the difference in prices.

One could argue that the coefficients on the recent privatization decision obtained at the estimations of columns 2 and 3 are rather high in absolute terms. Translated into euro units, this coefficient means that recent privatization reduces cross-subsidization by more than 40 euro cents per cubic meter, a quite dramatic change for residential users.<sup>(32)</sup> The inclusion of extra explanatory variables as the log of per-user consumption levels could contribute to check the robustness of such a high coefficient. In the instrumental variable regression of the OLS-instrumented case, the coefficient is indeed reduced, but not so much. The new coefficient is  $-59.5987$ , with a  $z$  statistic of  $-2.43$ . Although the coefficient is reduced, it still represents almost a 36 euro cent reduction of cross-subsidization. When we use total consumption levels instead, as extra variables, the regression increases the value of this main coefficient to  $-90.4393$  ( $z$  statistic is  $-2.20$ ). Finally, if I include both (log of) per-user consumption levels and total consumption levels (and I use all of the instruments), the coefficient of interest takes a value of  $-56.4445$ , with  $-2.35$  as the  $z$  statistic. I will keep the moderate value of approximately  $-55/-60$  (33–36 euro cent reduction) as the most reliable.

I briefly comment on the rest of the coefficients: the coefficient related to B3 is negative as predicted, although it is not significant in the majority of cases; the geographical coefficient is positive and significant—it accounts for unobservables that are specific for the half of Catalonia with internal river basins; the time dummy has no influence on the price difference, which is not very surprising. A further version of this model could try to include more variables explaining the difference of marginal costs, although a parsimonious model was called for in this case in the interests of simplicity. Concerning political variables, I observe that the mayor variable clearly does not have a significant effect on revealed regulator's preferences, while the social majority generally has a significant positive influence. This result does not contradict what is found in the literature on public service pricing/taxation. As an example, Bel and Miralles (2006) find that the level of service-specific deficit (cost minus revenues from the service-specific poll tax) in the solid-waste management service is better explained by the social-majority variable than by the mayor's ideology.<sup>(33)</sup>

Concerning the relative size of the industrial sector, the estimation reveals a positive effect, although it is significant in only one case. This suggests that the industrial interest group's effectiveness critically needs some special circumstances in which the regulator could argue in favor of a change in the status quo between consumer types. Otherwise, the substitution effect overrides the political support effect. By a feedback argument, this could explain why the industrial sector has recently lobbied for water-utilities privatization at the advent of the Water Framework Directive. In fact, when properly instrumented, the recent privatization decision explains

<sup>(32)</sup> A referee has brought my attention to the necessity of checking the robustness of this coefficient.

<sup>(33)</sup> In fact, a proper combination of both variables is what optimally explains the results.

more than the mere industrial-sector size. Finally, the coefficients on the probit-like equation explaining the privatization decision are all significant, and have the sign that they were expected to have had earlier in this section.

#### 4 Conclusions

One of the results observed in Miralles (2006) is that the industrial-sector water customers have lobbied for water service privatization from 1996 onwards. A natural extension of that work consists of assessing the extent to which they achieved reductions in the cross-subsidization status quo, which has usually been favorable to residential users. Consequently, this paper analyzes the effect of recent privatization on the difference between the marginal water price paid by the average industrial customer and the marginal price paid by the average residential user. The empirical analysis is based on a simple theoretical framework in which the regulator sets prices in order to maximize a weighted sum of surpluses and profits.

When I estimate the relevant regressions I find that the recent privatization variable is highly correlated with the error term. When correcting for it, I find that there is, in effect, a link between recent privatization and a reduction in the revealed regulator's preference for residential customers. Recent privatization of the service may have been taken as a chance to concede to industrial-sector demands for less cross-subsidization. In doing so, the regulator could avoid the political cost stemming from the alteration of the status quo between residential customers and industrial users. The change-of-regime context could have facilitated a discourse in favor of major changes in the tariff structure. By a feedback argument, this finding contributes to an explanation of why the industrial sector lobbied for the privatization of the water service in recent years.

Other explanatory variables are included in order to explain cross-subsidies between consumer types. Concerning sociopolitical variables, it is found that the social majorities, measured from the votes in national elections, contribute more to explain the variable of interest than merely the mayor's ideological 'side'. More insight in this field could be an interesting goal for further research. A proper combination of these two factors could better explain regulatory decisions. Also, a wider, longer database could contribute to confirm the major findings of this paper—this is also left for further research.

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