Semiparametric Estimation and Testing in a Model of Environmental Regulation with Adverse Selection[1]

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Abstract

We consider a principal-agent model of environmental regulation with adverse selection, where firms are regulated through contracts. We show how the model allows to recover information on structural cost parameters. We use a semiparametric method to estimate consistently such parameters without specifying the distribution of agents' types. We also show how to check for the specification of the econometric model, as well as auxiliary parametric assumptions, by means of specification tests based on nonparametric estimation. Results are used to analyze some economic issues related to environmental regulation.

JEL classification: C14 Semiparametric and Nonparametric Methods; D82 Asymmetric and Private Information.

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1 Introduction

The economics of information and incentives has experienced great developments in recent years. In particular, the design of contracting procedures has been the subject of a vast literature, see the monographs of Laffont and Tirole (1994) and Salanié (1997) for references. In the economic theory of contracts, the agent is characterized by a *private information* which determines his ultimate actions. Two possible sources of asymmetric information can be distinguished : (a) an unobservable action undertaken by the agent (effort in a production process, protection against risk,...); (b) an unknown characteristic of the agent (efficiency in terms of cost, willingness to pay for a given good,...). Case (b) is labelled *adverse selection* in the literature, and the agent's characteristic is referred to as his *type*. While the type of the agent is unknown to the principal, the latter nevertheless is assumed to have prior information before the contract is negociated, in terms of the statistical distribution of the type and other relevant characteristics of the agent. The challenge for the principal is to set up a contract scheme enforcing truthful revelation of the private information, thus allowing a second-best optimum to be attained for the economic variable of interest (production level, environmental externality, . . .).

In practice, contracts are largely used in domains as various as insurance, public regulation, industrial relationships, agricultural production or employment procedures. This suggests that asymmetric information can be present in many situations. Neglecting this issue is likely to lead to unreliable results. For instance, in a study of regulated water utilities, Wolak (1994) finds that explicitly modeling asymmetric information yields estimates of the production technology that are substantially different from the ones derived from a model ignoring the problem. However, estimation of structural models with adverse selection poses theoretical and practical problems. These empirical difficulties explain that much applied work is based on reduced-form models, see Chiappori and Salanié (2000) for references. Few econometric applications are based on structural models, see e.g. Auriol, Ivaldi and Kim (1999), Gagnepain and Ivaldi (1998), Ivaldi and Martimort (1994), Miravette (1997), Thomas (1995), Wolak (1994). They generally adopt a framework that requires parametric assumptions on the distribution of the agent's type. Such assumptions can be an important source of misspecification errors, whose impact is difficult to assess.

Our goal is to propose a flexible framework for estimating and testing a structural model of environmental regulation with adverse selection. This framework uses semiparametric methods for estimating consistently parameters of interest and for assessing the results by testing procedures. Specifically, the method is nonparametric with respect to the adverse selection parameter distribution. The advantages are twofold. First, from an estimation viewpoint, a semiparametric model allows to let the distribution of the agent's type unspecified, so that robust estimates of some structural parameters can be obtained. Moreover, the estimation does not require costly numerical algorithms to be used. Second, from a testing viewpoint, it is possible to check the semiparametric model through tests relying on nonparametric methods. As such tests are consistent against any alternative, they can detect misspecifications that are not uncovered by standard parametric tests. Furthermore, they also allow to test for maintained parametric assumptions that determine estimation of the structural parameters.

The paper is organized as follows. Section 2 gives an account of the problems related to structural estimation in models of adverse selection and presents different methods used in empirical work. Section 3 describes a simple model with adverse selection, where the principal is a local environment protection agency, and the agent is a polluting firm whose effluent emissions are to be reduced. In this case, the source of asymmetry lies in a private information parameter reflecting the efficiency of the firm in abatement activity. In Section 4, we present the econometric model jointly with the estimation and testing procedures. In Section 5, we apply our econometric framework to a sample of French industrials for the period 1987-1992 and we detail the economic implications of our findings. The Conclusion brings together some remarks on our approach and how it could be extended to more general setups. Appendices A to C gather technical details.

2 Estimation of structural adverse selection models : a non-exhaustive survey

Different problems arise for the econometrician when dealing with a structural adverse selection model. Indeed, many features of the model are unobservable. Private information is unobserved by both the principal and the econometrician, while prior information of the principal on the agent's type (in terms of statistical distribution) is also unknown to the econometrician. Other parameters of interest, such as the opportunity cost of public funds in public regulation, are also unknown. This means that identification of all structural parameters can be difficult, or even impossible, without imposing some restrictions on parameters. However, optimal contract solutions explicitely depend on such parameters, and crucially on the distribution of the private information. An approach used by most authors consists in assuming a parametric or semiparametric form for the type's distribution. This distribution should then be estimated jointly with other structural parameters. Hence, the resulting models are highly nonlinear with unobservable (latent) variables and estimation requires sophisticated and costly numerical procedures, such as direct numerical integration or simulation–based methods, see Laffont, Ossard and Vuong (1995) for an application to auctions.

We now detail selected empirical work from the literature, that are representative of the type of estimation methods proposed to deal with such models. In a purely methodological paper, Feinstein and Wolak (1991) explore the econometric implications of incentive compatible regulation. Their framework is used by Wolak (1994) in an empirical analysis of regulated water utilities in California. The key element to the model is asymmetric information on the quality of labour hired by water utilities. The estimation procedure consists in imposing semiparametric restrictions on the type's distribution and in computing the conditional expectation of variable cost with respect to its distribution. Estimation results reveal that overlooking the presence of asymmetric information produces biased parameter estimates for cost, and therefore for predicted returns to scale. Whereas under perfect information, water utilities operate close to the minimum efficient scale, there are significant decreasing returns to scale in the asymmetric information case. Using Vuong (1989) test for non-nested models, Wolak shows that the perfect information specification is rejected.

Ivaldi and Martimort (1994) consider the relationship between a firm and an energy input supplier. Because of oligopolistic price competition between energy suppliers, contracts are helpful in recovering suppliers' strategies. Asymmetric information is present because of heterogeneity among producers in their preferences for energy type. Empirical evidence from French data on dairy firms shows that the quadratic form of contracts originates from asymmetric information. The econometric model consists of a system of nonlinear equations with input price and quantities as endogenous variables. Assuming the private information parameter has a Beta distribution, the likelihood function is computed by a numerical integration technique. Estimation results reveal that the gas supplier tends to subsidize industrial dairies as more energy demanding industries.

Thomas (1995) develops a model of industrial pollution control with imperfect effluent emission tax and subsidized abatement activity. The basic assumption is that contracts between an environmental protection agency and polluting industrials act as a complementary policy to emission tax, which is suboptimal for administrative and price-control reasons. Asymmetric information is related to firm's efficiency in abatement of effluent emissions. A system of structural equations from the optimal contract first-order conditions is estimated by integrating out the private information parameter. This is achieved by using a Simulated GMM (Generalized Method of Moments) technique with an adequate choice of the importance sampling function. Estimation results reveal that the ratio of optimal (Pigouvian) over actual emission tax level is about two.

Auriol, Ivaldi and Kim (1999) discuss the role of advertising expenditures by regulated telephone companies in the U.S. The regulator problem is to maximize expected net consumer surplus without precise knowledge of producer cost, given that the net surplus is decreasing with the level of advertising. The inverse demand function for telephone and the firm cost equation are specified parametrically, using data on 67 U.S. companies from 1988 to 1996. Maximum likelihood estimates are obtained for the private information parameter and its distribution from the estimated cost function. These estimates are inserted in the first-order condition defining the optimal contract, which is estimated by G.M.M. The authors conclude that the optimal advertising level is about 30 percent of the actual level, and determine the resulting welfare loss as a fraction of average firm revenue.

These papers, as well as others that we do not detail here, impose different parametric restrictions on structural parameters. A first set of restriction is related to the agent's behavior through his cost or utility function for instance. A second type of restrictions concerns the type distribution, which is usually assumed to belong to a specific family. The third kind of restriction pertains to some parameters, such as the opportunity cost of public funds, which is set to an arbitrary value, see Auriol, Ivaldi and Kim (1999) and Wolak (1994). Our following analysis is based only upon restrictions of the first type, and avoids supplementary parametric assumptions on the type's distribution or auxiliary parameters.

3 Economic Modeling of Environmental Regulation

3.1 The Basic Abatement Model

Consider a firm whose production generates some effluent emission level B. The firm is equipped with an abatement plant whose abatement rate on B is denoted by δ. The outcoming (net) emission level from the abatement plant is therefore equal to $(1-\delta)B$. Environmental regulation is implemented by a local environment protection agency, which designs an emission charge t and grants subsidies T to support the firm's abatement activity. The emission tax is based on *actual* effluent emissions after abatement, i.e. $(1-\delta)B$. The subsidies are transfers covering part of the firm's operating and maintenance variable costs, but are not systematically granted, depending on the regulatory policy adopted. We assume that production and abatement activity are independent, which is justified in our application as we focus on *external* (end-of-pipe) abatement, and not clean technologies, for which production and abatement activities are technically entangled.[2] Hence, production output level and effluent emission B are fixed when considering abatement decisions. Therefore, the firm receives a positive amount $t\delta B$ from the regulator when abating the emission level B by $100 \times \delta$ percent, plus a transfer T. With c denoting the abatement cost, profit in the abatement activity reads

$$\Pi = t\delta B - c + T. \tag{1}$$

[2]Other justifications include the fact that the level of the pollution tax is low in practice compared to marginal profit from production.

We assume that c depends upon the emission inflow B, the pollution abatement rate δ and on an efficiency parameter θ as given in the following assumption.

Assumption 1 *$c = \theta C(B, \delta)$ with $\theta > 0, C'_B > 0, C'_\delta > 0, C''_{\delta\delta} > 0, C''_{B\delta} > 0$, where $C'_.$ and $C''_{..}$ denote first and second-order derivatives with respect to the arguments specified by the indices.*

The cost function is increasing in inflow of effluent and in abatement rate, and is convex in the abatement rate δ. The parameter θ is a (positive) scale parameter for cost. Since for a given level of emission inflow B and a given abatement rate δ, a firm with a higher parameter θ will have higher costs of abatement, θ is to be interpreted as an *inefficiency* parameter. The crucial assumption here is related to the separability of the cost function with respect to θ on the one hand and its other arguments on the other hand.

Consider the case where the firm only faces a unit emission tax t. We label this case the *Status-Quo* or *no regulation case*, because no contract exists to specify the degree of abatement to be achieved by the firm, i.e. $T = 0$. The firm will therefore select the abatement rate by maximizing profit with respect to δ, which yields $tB = \theta C'_\delta$. The Status-Quo abatement rate is then defined as

$$\delta^0 = [C'_\delta]^{-1} \left[\frac{tB}{\theta} \right]. \tag{2}$$

Because the abatement marginal cost is increasing in δ, the abatement rate is increasing in the unit emission tax, the emission inflow, but decreasing in the inefficiency parameter θ.

In the theory of environmental regulation, the so–called Pigouvian tax, which is equal to marginal utility of consumers for abatement, achieves the *socially-optimal* rate of abatement. Where firms are faced with the Pigouvian tax, they fully internalize the social cost of environmental damages caused by their pollution. This is known as the *Polluter-Payer* principle. In practice, the emission charge can be restricted for practical and institutional reasons, see Thomas (1995). Furthermore, when consumers' preferences towards pollution are not known with sufficient accuracy, the regulator may not be able to design the proper Pigouvian tax scheme, see Baumol and Oates (1988). As a result, firms facing an inadequate emission charge will not find it profitable to abate at a "socially-acceptable" rate.

The problem is complicated further because uniformity of the emission tax does not allow for optimal policies to be achieved in practice. This is because distortions are likely to be important when considering heterogenous polluters, characterized by different inefficiency parameters. Theoretically, this could be overcome by letting the tax t depending on θ. However, in most real–world environmental policies, the emission charge is in fact uniform, mainly for ease of implementation and on equity grounds. Hence, it is often not possible for the regulator to implement a personalized, firm–specific emission charge.

It is then clear that more flexible policies are called for. A complementary environmental policy instrument which is often used in practice is a contract between the local regulator and the firm. According to this contract, the firm accepts to abate at a given rate δ, while receiving a subsidy indexed on δ. The contract-based regulation policy allows for a case-by-case determination of abatement rate consistent with standard regulatory and juridical procedures. It is thus interesting as a complementary or alternative policy to the uniform emission charge regulation. As we will see below, the actual performance of contract schemes crucially depends on the ability for the regulator to observe firm's characteristics that affect the abatement activity.

3.2 Regulation under Perfect Information

In the contract-based regulation, the firm commits to abate at a specific rate δ while receiving a transfer T. The regulator is assumed to maximize total surplus, i.e. for consumers and the firm, with respect to abatement level. The consumers' utility function for abated pollution is denoted $W(\delta B)$, with $W' > 0$, $W'' \leq 0$. Both the subsidy T and the amount paid to the firm for abatement, $tB\delta$, must be paid by consumers through some redistribution (fiscal) mechanism. The regulator preferences are summarized in a parameter σ, used as a weight in the surplus function. A high value for this preference parameter indicates that the regulator favors consumers more than the firm. Assuming that consumers' surplus and regulator's budget can be aggregated, total surplus writes

$$
\begin{aligned}
&\sigma\left[W(B\delta) - tB\delta - T\right] + (1-\sigma)\Pi(\theta,\delta,T) \\
&= \sigma\left[W(B\delta) - \theta C(B,\delta)\right] + (1-2\sigma)\Pi(\theta,\delta,T).
\end{aligned}
$$

using profit's definition (1). Following Baron (1989), the preference weight must be such that $0.5 \leq \sigma \leq 1$ for the problem to be consistent. Indeed, a somewhat different exposition of the problem shows that values of σ less than 0.5 correspond to a *negative* opportunity cost of public funds.

An optimal contract between the regulator and the firm consists in a pair $(\delta(\theta), T(\theta))$, which maximizes total surplus. For the firm to accept to participate in the contract relationship, its profit under regulation must be greater than its Status-Quo profit. This condition is labelled as *Individual Rationality* in the literature. Under perfect information, i.e. when the regulator knows the value of θ, the problem is simplified by the fact that the principal is able, through the transfer, to exactly equate both regulation and Status-Quo profits. Hence, the second term in total surplus can be omitted. The solution δ^* is obtained by maximizing total surplus with respect to the abatement rate, which yields

$$W'(B\delta^*)B = \theta C'_\delta.$$

This is the standard result of Pigouvian taxation: the non-uniform rewarding scheme of contracts allows to attain the *First-Best* solution, which equates marginal utility and marginal cost of abatement.

As argued above, it is likely that emission charge does not accurately reflect consumers' utility for abatement. To formally take this into account, we make the following assumption.

Assumption 2 *Marginal utility for abatement is proportional to the emission charge, i.e.*

$$W'(\cdot) = t/\varepsilon, \text{ with } 0 < \varepsilon \leq 1.$$

The parameter ε thus summarizes the imperfection in the uniform emission tax scheme. With this assumption, the First-Best abatement rate is

$$\delta^* = [C'_\delta]^{-1}\left[\frac{tB}{\varepsilon\theta}\right]. \tag{3}$$

It is easy to see that since $\varepsilon \leq 1$, the abatement rate under regulation with perfect information is always greater than the Status-Quo abatement rate δ^0, so that regulation is beneficial in terms of environmental conservation.

3.3 Regulation under Asymmetric Information

We now consider the more realistic case where the principal does not observe the agent's inefficiency parameter θ. In theory, the contracting procedure is modeled as follows. First, the firm is asked to report its type (or private information), i.e. its inefficiency parameter. Based on the reported $\tilde{\theta}$, the regulator then proposes a contract, i.e. a pair $(\delta(\tilde{\theta}), T(\tilde{\theta}))$, that the firm can accept or not. Nevertheless, it is not necessary in practice to require the agent to reveal his type directly. The regulator offers a menu of contracts to the agent who, assuming the contracting scheme is properly designed, selects the contract corresponding to his type.

Obviously, the contracting scheme has to satisfy the Individual Rationality constraint that profit under regulation is not less than the Status-Quo profit, i.e. $\Pi(\theta, \delta^*(\theta), T(\theta)) \geq \Pi(\theta, \delta^0, 0)$. Moreover, another constraint is relative to the revelating property of the contract scheme. The regulator is willing to enforce truthful reporting of θ by the firm, in order to avoid strategic behavior. Indeed, as the contract is indexed on the reported parameter $\tilde{\theta}$, an efficient firm, i.e. a firm with a low θ, may instead report itself as an inefficient one, i.e. with a high θ. In so doing, the firm may be assigned a reasonable objective in terms of abatement and makes a greater profit than if telling the truth. The constraint associated to truthful revelation is denoted *Incentive Compatibility* in the literature. It states that the firm must be better off reporting its true type, as it would be reporting any other value. A major difference with the First-Best solution is that the regulator now has to grant an *information rent* to the firm in return for its truthful report. Such a rent is granted through the transfer $T(\cdot)$. Because of the cost of public funds, consumers experience a financial burden related to information rents.

Therefore, the regulator is not able to implement the First-Best abatement rate, and in particular is unable to determine the transfer which would achieve equality between profits under Status-Quo and regulation. He has nevertheless prior information on the firm's characteristic, which enables him to maximize total surplus *ex ante*, over the definition domain for θ. Such information is available to the principal through past contracts with similar firms, or technical data on abatement activity on a sector-by-sector basis. Prior information to the regulator is traditionally represented by a probability distribution function $F(\theta)$ with associated density function $f(\theta)$, defined on a domain $[\underline{\theta}, \overline{\theta}]$. Total expected surplus then writes

$$\int_{\underline{\theta}}^{\overline{\theta}} \{\sigma [W(B\delta(\theta)) - \theta C(B, \delta(\theta))] + (1-2\sigma)\Pi(\theta, \delta(\theta), T(\theta))\} f(\theta) d\theta. \tag{4}$$

The menu of optimal contracts is chosen so as to maximize (4), taking into account the constraints

$$\Pi(\theta, \delta(\theta), T(\theta)) \geq \Pi(\theta, \delta(\tilde{\theta}), T(\tilde{\theta})) \quad \forall\, \theta, \tilde{\theta} \quad \text{(Incentive Compatibility)}, \tag{5}$$

$$\Pi(\theta, \delta(\theta), T(\theta)) \geq \Pi(\theta, \delta^0, 0) \quad \forall\, \theta \quad \text{(Individual Rationality)}, \tag{6}$$

where θ is the true parameter value and $\tilde{\theta}$ is the report by the firm of its type. Because our Assumption 1 implies $C'_\delta > 0$, the so-called single-crossing (Spence-Mirlees) condition holds, as is usually assumed in the principal-agent theory. We thus obtain a separating equilibrium. As shown in Appendix A, the equilibrium abatement rate under asymmetric information is given by

$$\delta^{**} = \begin{cases} [C'_\delta]^{-1} \left[\frac{tB}{\varepsilon H(\theta)}\right] & \text{for } \theta < \theta_c, \\ [C'_\delta]^{-1} \left[\frac{tB}{\theta}\right] & \text{for } \theta \geq \theta_c, \end{cases} \tag{7}$$

where

$$H(\theta) = \left[\theta + \frac{(2\sigma - 1)}{\sigma} \frac{F(\theta)}{f(\theta)}\right], \tag{8}$$

and θ_c is solution of the equation $\theta_c = \varepsilon H(\theta_c)$, so that δ^{**} is continuous at θ_c. Figure 1 presents the different solutions for the abatement rate in function of θ. It is easily seen that the abatement rate solutions can be ranked as follows: $\delta_0(\theta) \leq \delta^{**}(\theta) \leq \delta^*(\theta), \quad \forall\, \theta$. That is, the contracting procedure allows to attain higher abatement rates than in the Status-Quo, but lower than the First-Best ones.

Only firms with inefficiency parameter *less than* θ_c are regulated by means of contracts. On the contrary, firms above the threshold value are not regulated and are left at the Status-Quo level, namely they will be characterized by the abatement rate δ^0. This comes from two competing effects. On the one hand, it is socially more profitable to regulate efficient firms so as to promote overall emission reduction. On the other hand, it is more costly to regulate efficient firms since the information rent is decreasing in θ, see Appendix A. The principal therefore concentrates only upon a fringe of firms below the threshold value θ_c, so as to reduce overall information rents. The threshold value crucially depends upon the discrepancy between the Pigouvian tax level and the actual emission charge as measured by ε, and upon the preference parameter σ.

4 Econometric Modeling

4.1 Model Specification and Estimation

Our goal is to estimate parameters of the abatement cost structure from the first-order condition of the principal-agent model. These depends upon the form of the marginal cost with respect to abatement rate. Therefore, to obtain a tractable econometric model, we make the following assumption on the abatement cost function, see McConnell and Schwarz (1992) for a similar specification and the rationale behind it.

Assumption 3 *The abatement cost function takes the Cobb-Douglas form*

$$c = \theta C(B, \delta) = \theta B^{\alpha_0} \delta^{\alpha_1}, \text{ with } \alpha_0 > 0, \ \alpha_1 > 1.$$

We can then derive explicit forms of the abatement rates solutions of (2), (3) and (7). Under Assumption 3, the Status-Quo solution is

$$\log \delta^0 = \frac{1}{\alpha_1 - 1} \left[-\log \alpha_1 + \log t - (\alpha_0 - 1) \log B - \log \theta \right].$$

Similarly, the First-Best abatement rate is

$$\log \delta^* = \frac{1}{\alpha_1 - 1} \left[-\log \varepsilon \alpha_1 + \log t - (\alpha_0 - 1) \log B - \log \theta \right],$$

while the abatement rate under asymmetric information when contracting takes place is

$$\log \delta^{**} = \frac{1}{\alpha_1 - 1} \left[-\log \varepsilon \alpha_1 + \log t - (\alpha_0 - 1) \log B - \log H(\theta) \right] \qquad \text{for } \theta \leq \theta_c. \quad (9)$$

In what follows, we focus on the latter equation. Indeed, we are primarily interested in estimation of (9), because our real case concerns contracts between firms and regulation agencies where asymmetric information about abatement costs is present. Though, the rationale that we follow equally apply to the estimation and testing of the econometric models derived in the Status-Quo and First-Best situations.

Equation (9) determines the abatement rate through an equation of the form

$$\log \delta = \beta + X'\lambda + V,$$

where $X = (\log t, \log B)$ are explanatory variables, λ are parameters derived from structural parameters of the economic model, and $V = \log H(\theta)/(1-\alpha_1)$ is a random unobservable component. It is quite remarkable that the theoretical economic model gives directly rise to an econometric model with random elements, so that the econometrician does not need to artificially supplement the economic model with an error component. Though, the element V cannot be identified with a standard residual for several reasons. First it is likely not to be centered, so that estimation of the constant term will not convey any information on the structural parameters. Second, in practice, the distribution of the firms' type are likely to differ upon some characteristics of the firm, such as the industry sector in which it operates. If relevant, this information is to be used by the regulator and should be taken into account by the econometrician as well. It is then necessary to introduce these observable characteristics as supplementary explanatory variables in the model.[3] However, there is no clear way to formulate the relationship linking θ and firms' characteristics Z. Hence we describe this relationship through a nonparametric function $g(\cdot)$, so that

$$V = g(Z) + U, \;\; E\left[U|Z\right] = 0.$$

This enables us to acknowledge the influence of the characteristics Z on the first moment of $\log H(\cdot)$ and leads us to consider an econometric model of the form

$$\log \delta = X'\lambda + g(Z) + U, \tag{10}$$

where $g(\cdot)$ is some unknown function of the characteristics Z that includes the constant. Because U is related to θ, which is predetermined with respect to both effluent emission and emission charge, it has null conditional expectation with respect to all explanatory variables entering the model. In our framework, there is no need for supplementary distributional assumptions on the error term U.[4]

[3]This is akin to considering that the prior distribution used by the principal is conditional to observable characteristics.

[4]Obviously, there may be potential influence of Z on other moments than the first. If the econometrician has some a priori information, this influence may be included as well in a parametric or nonparametric

Our final model is a partial linear regression model, and different estimation methods allow to estimate λ and $g(\cdot)$. In particular, parameters of the linear part can be $\sqrt{n}$-consistently estimated by some semiparametric estimation technique. Specifically, we will use Robinson's (1988) method, see Appendix B for a presentation of the method. These estimates can in turn be used to recover estimates of cost parameters α_0 and α_1. Such estimates will be robust with respect to the unknown distribution of the agents' type, while accounting for characteristics that could influence the types' distribution. They are obtained without actually observing cost data, but only abatement decisions of contracting firms. The basic requirement of the method is that the effects of the agents' type and of the other variables in the cost function can be separated.

Previous work dealing with estimation of asymmetric information structural models used to specify distributional assumptions for the parameter θ conditional upon characteristics Z. This allows to recover estimates not only of the cost function's parameters but also of the conditional distribution of θ given Z. However, in the absence of specific information, the types' distribution is chosen somewhat arbitrarily, and robustness of the results to such an assumption is far from ensured. In contrast, our estimation method allows to characterize the determinants of the cost structure without the need for specifying the types' distribution.

4.2 Specification Checks

The validity of the estimation method heavily relies on our maintained Assumptions 2 and 3. In particular, the separation between the type θ and other factors in the profit function is central for our analysis. This hypothesis, together with the Cobb-Douglas specification of the cost function, allows estimation of structural parameters of interest. Therefore, it is crucial to check whether the specification of the economic model is acceptable in view of the evidence provided by the data.

Several procedures for testing a parametric or semiparametric specification against a nonparametric one exist in the literature. In our application, we will use the one developped by Zheng (1996) for a parametric model and further developped by Fan and Li

fashion, so as to get more efficient estimators of the basic structural parameters. We will not pursue this here.

(1996) for a semiparametric model. The procedure is built to test the whole specification of the model, through the null hypothesis $H_0 : E[U|X, Z] = 0$, see Appendix C for details. Another specification check can be entertained on our econometric model. Indeed, linearity of $\log \delta$ with respect to $\log t$ and $\log B$ is entailed by our assumptions. A specification check for this linearity can be implemented by extending the testing framework developped in Delgado, Dominguez and Lavergne (1998) for testing conditional moment restrictions to the case where the null model is partially linear. This extension is quite straightforward, see Appendix C for some explanations.[5] However, such a test provides an unexpected application of the specification testing framework, where the model is a regression model, but we are interested in a conditional moment restriction which involves only some of the regressors, that is $E[U|X] = 0$. We label it a specification test of the *parametric part* of the model.

An advantage of the latter test is that it considers a moment condition that depends on fewer variables than the former full specification test. This is likely to be valuable because the behavior of tests for conditional moment restrictions based on nonparametric estimation can be quite sensitive to the number of conditioning variables. Hence, the asymptotic normal approximation may not provide an adequate one for a small sample size or when the number of conditioning variables is large with respect to the sample size. Indeed, the test statistic asymptotically behaves like a centered and rescaled chi-square with degrees of freedom converging to infinity and accordingly the finite sample distribution is typically right-skewed. Hence, in specification checks, it may be misleading to use p-values coming from the asymptotic normal approximation. A wild bootstrap procedure, such as studied by Härdle and Mammen (1993) and Li and Wang (1995), can be applied to compute more accurate p-values. Unfortunately, we cannot apply wild bootstrap in testing the specification of the parametric part of the model, as we are testing for the nullity of a conditional expectation involving only part of the explanatory variables, i.e. only X. In this case, a bootstrap procedure would need to generate values of Z that mimic the dependence with the other variables.

[5]Delgado, Dominguez and Lavergne (1998) deal with testing conditional moment restrictions in a parametric model. The formal analysis of the conditional moment restriction test for a semiparametric regression model follows similar lines, see Fan and Li (1996) for specification testing of a partial linear regression model.

5 Empirical Application

5.1 Data Description

We use plant–level data on abatement activity of French firms. The plants are located in the French hydrographical basin corresponding to the Seine–Normandie (Paris and North of France) local water agency. The number of observations is 213 and the sample period is 1987-1992. Contracts are recorded between industrials and the water agency, concerning external abatement plants only. This is because they do not modify the production process and allow to recover abatement variables, which would not be possible with internal abatement. There are some cases in which contracts cover several successive operations. Roughly speaking, the typical case consists in a first period where the contract concerns fixed costs of a new abatement plant, and a second period where the contract concerns variable costs of the settled plant. Abatement rates are then computed after complete setting up of the plant, taking into account the one–year delay for the equipment to become fully operative. Therefore, each firm is represented only once in the sample.[6]

Effluent emission data are available on five categories of pollutants: Biological Oxygen Demand (BOD), Total Suspended Solids, Nitrogen, Phosphates, and Inhibitory Matters. We choose BOD as the pollution index, as it is good indicator for overall pollution, accounting for more than 75% of total emission fees.[7] The pollution inflow B is then defined as the *gross* (i.e. before treatment) emission level of BOD, in kg per day. In the following, the abatement rate δ and the unit emission charge t also correspond to BOD. Table 1 presents descriptive statistics for variables in the sample. Variability in the unit emission charge originates both from yearly variation and from regional differences in the design of the emission fees.[8]

There exist three possible steps in wastewater abatement. The first and secondary steps,

[6]An implicit assumption is that there is no technical progress in the considered period. There is evidence that the abatement technology has not evolved much in the considered period.

[7]BOD is a conventional, degradable pollutant defined as the quantity of oxygen absorbed by the effluent, measured on a 5-day period at a temperature of 20^oC. BOD is a good measure of microbiological activity, particularly when the effluent is severely polluted. See McConnell and Schwarz (1992) for details.

[8]On the period considered, emission charges designed by local Water Agencies were fairly stable before the 1992 French Water Act induced a significant increase in unit emission fees from 1992 to 1996.

denoted primary treatment, deals with organic matters essentially (sediment, suspended solids, ...), while advanced treatment is required when effluent emissions are more toxic (nitrate, phosphorous, ...). A firm will be more efficient in the abatement activity when effluents can be abated with primary treatment only. Hence, we select an toxicity index of effluent emissions to explain the inefficiency parameter. This index is the part of toxic effluents in total emissions before treatment, and is denoted $PART$, with $PART$ is in $[0, 1]$. A value of that index close to zero indicates that effluent emissions are mostly organic, and therefore require limited additional treatment. On the contrary, when organic effluents come jointly with other more toxic pollutants, i.e. when $PART$ moves away from zero, pollution abatement requires more know–how and ability.

It is also likely that firms operating in different industries have different abilities in the abatement activity in a way that may not be directly captured by the toxicity index $PART$. Then, we also consider the industry sector of the firm (SIC) as potentially related to the inefficiency parameter. The definitions of the sectors are the following, with the number of firms in each sector in parentheses: Food and drinks (47 firms); Dairy products (25 firms); Chemicals (16 firms); Iron and steel (94 firms); Paper and wood (17 firms); Others (14 firms). Due to data limitation, we introduce corresponding dummy variables in an additive way in $g(\cdot)$.

5.2 Results

We estimate the econometric model (10) by Robinson's (1988) method. This method requires a bandwidth that asymptotically undersmoothes with respect to the theoretical optimal bandwidth in nonparametric regression estimation, which usually is of order $n^{-1/5}$ for an unidimensional Z. Thus, we choose bandwidth parameters proportional to $n^{-0.3}$. Further analysis (whose results are not reported) shows that our main results are not very sensitive to variation of this parameter. Table 2 reports our estimation results. In computing standard errors, we have taken into account possible heteroskedasticity by Eicker's (1963) method as suggested by Robinson (1988). The parameter related to $\log t$ is highly significant, while the one for $\log B$ is close to zero and non significant. However, when recovering structural parameters α_0 and α_1 from these estimates, we obtain estimates that are highly significant. Although we have not imposed restrictions on parameters ensuring that cost is convex in δ and increasing in B, our estimates fulfil these restrictions. The parameter α_0 is not significantly different from one, whereas the estimated α_1 is about 2.46 with standard deviation 0.42. Parameters related to the different industries appear to be insignificant, with the exception of the ones attached to the Iron and Steel industry, which is negative, indicating that firms in this sector are relatively less efficient than firms operating in other industries. The fact that the Iron and Steel industry seems to be among the less efficient ones in abatement activity was already noted by Thomas (1995).

We also perform specification tests on the semiparametric model. For checking the parametric part of the model, we apply the testing procedure detailled in Section 3.2. We compute an individual bandwidth for each variable $\log t$ and $\log B$. The choice of the bandwidth parameters uses the rule-of-thumb $h = a\ 0.79\ I\ n^{-1/5}$, where I is the interquartile range of the variable, see Härdle (1993). In order to investigate the sensitivity of the test to the choice of the bandwidths, we let the multiplicative factor a vary from 0.5 to 1.5. The test statistics in Table 3 all exhibit negative values. Because asymptotically negative values of the test statistic occur under the null only, the test is one-sided and p-values computed from the asymptotic normal approximation are all greater than 0.83. This gives strong evidence in favor of the parametric assumptions adopted.

Considering now checking for the full model specification, we apply the testing procedure detailled in Section 3.3. We use a similar method for the choice of the bandwidth parameters, now including the variable $PART$ in addition to $\log t$ and $\log B$. We do not account for the dummy variables corresponding to different industries, because most coefficients are not significant. Values of the test statistic are negative or slightly positive, and p-values based on the asymptotic approximation are always greater than 0.36. This indicates that our semiparametric model is a valid candidate for modeling the abatement equation.

Finally, Figure 2 shows the estimated function $g(\cdot)$ from the semiparametric model with a bandwidth proportional to $n^{-1/5}$. The curve is decreasing when $PART$ is less than 0.8, indicating that toxicity of effluent adversely affects efficiency in abatement. This effect is much stronger when the toxicity index is between 0.6 and 0.8. Though, after the threshold 0.8, efficiency is increased, reflecting an actual improvement in efficiency due to the more homogeneous nature of effluent.

5.3 Economic Implications

Important implications for economic and environmental policies can be drawn from our empirical results. First, viewing abatement as an industrial activity in its own right, the cost structure can be used to evaluate returns to scale for this activity. Second, estimates of marginal abatement costs are important as guidelines for cost subsidies in many countries, see e.g. McConnell and Schwarz (1992). Moreover, assessing the sensitivity of abatement rates to effluent emissions and effluent tax schemes is helpful in designing environmental policies.

5.3.1 Returns to Scale in Abatement Activity

A first useful concept is the *elasticity of scale* in terms of abatement rate, defined as the ratio of average over marginal cost, see e.g. Chambers (1988). The scale elasticity is helpful in measuring the marginal increase in abatement cost when the abatement rate increases, for the same emission inflow. There are increasing, constant, or decreasing returns to scale if this elasticity if less than, equal to, or greater than 1. In our case, the Cobb-Douglas specification for the abatement cost function leads to a scale elasticity equal to $1/\alpha_1$, which is always less than 1 under Assumption 3. Our results lead to an estimated elasticity of 0.4058, with a standard deviation of 0.0684. This indicates that marginal cost of abatement is about 2.5 times the average cost for the firms in the sample. Though, our computations do not account for the high fixed cost of construction of the abatement plant.

Another measure deals with elasticity with respect to emission inflow. In this case, we investigate the potential economies of scale when the level of input to an abatement plant is increased, for the same abatement rate. There are constant (respectively increasing or decreasing) returns to scale in the level of emission inflow, if for a proportional increase Δ in input B to the abatement plant, we have $\Delta C(B,\delta) = C(\Delta B,\delta)$ (respectively $\leq$ or $\geq$). With the Cobb-Douglas cost function, constant returns to scale with respect to B is equivalent to $\alpha_0 = 1$. From our estimation results, we cannot reject the hypothesis of constant returns to scale. In other terms, there is no comparative gain from integrating emission inflows of different production plants, and treating the aggregate inflow in a single abatement plant.

5.3.2 Implications for environmental policies

Obtaining consistent estimates of marginal abatement cost is central for implementing environmental policies, for two reasons. In order to reach a socially acceptable level of pollution, the regulator must compare the marginal utility for pollution abatement with the marginal cost of abatement for industrials. Consequently, it is valuable to estimate marginal costs with sufficient accuracy for assessing the impact of different regulatory schemes on overall pollution abatement. Besides, in many instances, subsidies for fixed or variable cost of abatement are indexed on average marginal cost of abatement, in selected industries or in the economy as a whole. A reliable estimate of such cost obtained from individual data can thus serve as a guideline for such subsidy policies.

In a situation where there exists a significant discrepancy between the optimal and the actual emission tax, the sensitivity of industrials to the tax is a crucial parameter. The elasticity of effluent emissions to the unit emission charge allows to predict the performance of a tax-driven environmental policy in terms of reduced pollution. It also indicates the gains to the regulator from such a policy in monetary terms. In the Status-Quo case, the elasticity of abatement rate to unit emission tax is simply

$$\frac{\partial \log \delta}{\partial \log t} = \frac{1}{\alpha_1 - 1},$$

which is positive as $\alpha_1 > 1$. In the asymmetric information case however, the impact of a change in the unit emission tax in the abatement rate depends on whether the firm is regulated through a contract or not. The above "direct" effect only appears for firms that are not contracting, as for regulated firms, the subsidy constitutes a supplementary instrument that cancels any effect of the tax scheme. Hence, we can formally write

$$E\left[\frac{\partial \log \delta}{\partial \log t}\right] = \int_{\theta_c}^{\overline{\theta}} \frac{\partial \log \delta^0(\theta)}{\partial \log t} f(\theta)\, d\theta = (\alpha_1 - 1)^{-1} \left[1 - F(\theta_c)\right].$$

Therefore, $(\alpha_1 - 1)^{-1}$ appears as an upper bound on the relative effect of the tax on the abatement rate. An estimate of this upper bound is directly provided by our results and specifically equals 0.68 for our sample. Interestingly, a rough estimate of $1 - F(\theta_c)$ can be obtained by the regulation agency as the percentage of firms that do not enter the contract-based regulation.

Our results also allow to evaluate the sensitivity of abatement rate to emission inflow. In our case, the elasticity of δ with respect to B is given by

$$\frac{\partial \log \delta}{\partial \log B} = -\frac{(\alpha_0 - 1)}{\alpha_1 - 1}.$$

Note that this elasticity is negative in the case of decreasing returns to scale with respect to emission inflow. From our results, it appears that this elasticity, which corresponds to the coefficient of $\log B$ in our semiparametric model, is not significantly different from zero. One may wonder whether such a conclusion is reached because of our initial assumption that production and abatement decisions are sequential. That is, the production level may not be exogenous in our econometric model. To check for this assumption, we entertain estimation of our semiparametric model (10) with possibly endogeneous B.[9] To instrument B, we consider the level of capital stock invested in the abatement plant, which is highly correlated with the gross emission level. Estimation results are given in Table 2. They reveal that the parameter related to B has not significantly changed relative to the original model. A Hausman-type test formally confirms this, with a p-value of 0.47. Moreover, estimation results show that none of our main findings are affected.

6 Conclusion

The objective of this paper is to propose a semiparametric framework for estimating and testing a model of adverse selection for environmental regulation. The estimation method is computationally simple. It allows to recover structural parameters of the model in a manner that is robust to the distribution of the adverse selection parameter. These estimates enable in turn to analyse several economic issues. Specification checks based on nonparametric methods can be used to validate the auxiliary parametric assumptions used for estimation purposes.

While our framework is applied to a specific problem of environmental regulation, we believe that it can prove useful in other types of problems where adverse selection is present. Our work focuses on how existing techniques can be adequately used to deal with

[9]Robinson (1998) discusses such a generalization of his method in Section 7.2. A formal analysis of this method follows the lines of his proofs.

a structural equation derived from a principal-agent problem. The approach is motivated by the aim to obtain robust estimates of some parameters of interest independently from any particular specification of the type's distribution. It then precludes estimation of the distribution of θ as well as of the principal's preference parameter σ, but avoid strong restrictions on these parameters. Our model only relies upon assumptions that allow to obtain a tractable econometric model. The first restriction implied by Assumption 1 is the separability of the marginal cost function with respect to θ on the one hand and its other arguments on the other hand. This could be weakened, as soon as there is a known transformation of marginal cost that is separable. Such a separability is likely to be necessary to obtain a estimable econometric model. Otherwise, one should introduce other assumptions such as distributional ones on the adverse selection parameter. The second restriction implicit in Assumption 1 is that the adverse selection parameter simply enters in the cost function as θ itself only. But this is not restrictive if one lets its distribution unrestricted. Assumption 3 is of a parametric form. The Cobb-Douglas specification allows to write the econometric model as a simple regression equation, so that existing procedures can be easily applied or extended. Other specifications can introduce nonlinearities in the regression model that require a case-by-case treatment. This is outside the scope of this paper, but some hints can be found in the recent econometric literature. Robinson (1988) provides useful suggestions for extending his original method to the case of nonlinearities in explanatory variables. Also, nonlinearities could appear in the right-side of the equation. In that case, estimations methods developped for single-index models may be used to recover in a first step parameters related to explanatory variables, and the remaining parameters could be estimated in a second step, on such methods see Horowitz (1998). Depending on the final form of the econometric model, our testing approach may need customization to check for the auxiliary parametric assumptions that determine estimation of the parameters.

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Appendix A. Derivation of the Asymmetric Information Solution

We briefly sketch the resolution of the principal's problem, using standard techniques in the literature on regulation under asymmetric information. The interested reader may consult Baron (1989) or Laffont and Tirole (1994) for more details. Let us denote $\Pi(\theta,\tilde{\theta}) = \Pi(\theta,\delta(\tilde{\theta}),T(\tilde{\theta}))$ and $\Pi^0(\theta) = \Pi(\theta,\delta^0,0)$. We consider first the Incentive Compatibility constraint, i.e.

$$\Pi(\theta,\theta) \geq \Pi(\theta,\tilde{\theta}) \qquad \forall\,\theta,\tilde{\theta}.$$

This is equivalent to $\frac{\partial \Pi(\theta,\tilde{\theta})}{\partial \tilde{\theta}} = 0$, $\forall \tilde{\theta} = \theta$. Differentiating profit totally at $\tilde{\theta} = \theta$ yields $\dot{\Pi}(\theta,\theta) = \frac{\partial \Pi(\theta,\tilde{\theta})}{\partial \theta} = -C(B,\delta) \leq 0$. Hence, when the firm maximizes profit, total profit is decreasing in θ. Because the Spence-Mirlees (single-crossing) condition $C'_\delta > 0$ follows from Assumption 1, the second-order condition is verified as soon as

$$\frac{d\delta^{**}}{d\theta} \leq 0.$$

This in turn implies that $\frac{dT^{**}}{d\theta} \leq 0$.

Let both profits $\Pi(\theta,\theta)$ and $\Pi^0(\theta)$ coincide at some value of θ, denoted θ_c, which may coincide with $\underline{\theta}$ or $\overline{\theta}$. Taking into account the Individual Rationality constraint, it follows that firms above this threshold are left at the Status-Quo abatement rate.

Let us now solve the problem. Integrating by parts the expectation of profit yields

$$\int_{\underline{\theta}}^{\theta_c} \Pi(\theta,\theta) f(\theta)\,d\theta = \Pi(\theta_c)F(\theta_c) - \int_{\underline{\theta}}^{\theta_c} \dot{\Pi}(\theta,\theta)F(\theta)d\theta = \Pi^0(\theta_c)F(\theta_c) + \int_{\underline{\theta}}^{\theta_c} C(B,\delta(\theta))F(\theta)d\theta.$$

Total expected surplus then reads

$$\int_{\underline{\theta}}^{\theta_c} \left\{\sigma\left[W(B\delta) - \theta C(B,\delta)\right] + (1-2\sigma)\frac{F(\theta)}{f(\theta)}C(B,\delta(\theta))\right\} f(\theta)d\theta + (1-2\sigma)\Pi^0(\theta_c)F(\theta_c).$$

Maximizing with respect to δ yields

$$W'B - \theta C'_\delta + \frac{(1-2\sigma)}{\sigma}\frac{F(\theta)}{f(\theta)}C'_\delta = 0.$$

Hence the Second-Best (asymmetric information) solution for the abatement rate is

$$\delta^{**} = \left[C'_\delta\right]^{-1}\left\{\frac{tB}{\varepsilon}\left[\theta + \frac{(2\sigma-1)}{\sigma}\frac{F(\theta)}{f(\theta)}\right]^{-1}\right\} = \left[C'_\delta\right]^{-1}\left\{\frac{tB}{\varepsilon H(\theta)}\right\}.$$

It is easily checked that under Assumption 1 the second-order condition holds as soon as $W'' \leq 0$. Also the solution is decreasing in θ if in addition

$$\frac{d(F/f)}{d\theta} \geq -1.$$

This equilibrium solution under asymmetric information is valid for firms with characteristics in the interval $[\underline{\theta}, \theta_c]$. Thus, θ_c is the value of θ such that $\delta^{**} = \delta^0$, i.e. θ_c is solution to $\theta_c = \varepsilon H(\theta_c)$. In the polar case where $\varepsilon = 1$ (no distortion), $\theta_c = \underline{\theta}$, provided $f(\underline{\theta}) \neq 0$.

Appendix B. Estimation of a partially linear regression model by Robinson's (1988) method

Consider the following partially linear regression model

$$Y = \beta + X'\lambda + g(Z) + U, \qquad E\left[U|X, Z\right] = 0 \ . \tag{11}$$

By taking conditional expectation with respect to Z on both sides, we get

$$g(Z) = E[Y|Z] - E[X'|Z]\lambda.$$

By difference, we end up with

$$Y - E[Y|Z] = [X - E(X|Z)]'\,\lambda + U. \tag{12}$$

The estimation procedure consists in inserting nonparametric (kernel) estimates $E_n[Y|Z]$ and $E_n[X|Z]$ in place of the unknown conditional expectations and estimating λ by a standard no-intercept OLS rule. The first-step nonparametric estimation requires a bandwidth that asymptotically undersmoothes with respect to the theoretical optimal bandwidth in nonparametric regression estimation, which usually is of order $n^{-1/5}$ for an unidimensional Z.

In the case of i.i.d. observations, Robinson shows that the resulting parameter estimate $\widehat{\lambda}$ is $\sqrt{n}$-consistent and asymptotically normal. More formally, if U is independent of (X, Z) and homoscedastic with variance σ^2, then under technical smmothness assumptions on different functions,

$$\sqrt{n}\left(\widehat{\lambda} - \lambda\right) \xrightarrow{d} N\left(0, \sigma^2\Phi^{-1}\right) \ ,$$

if and only if

$$\Phi = E\left[\left(X - E\left(X|Z\right)\right)\left(X - E\left(X|Z\right)\right)'\right] \text{ is positive semidefinite.}$$

Moreover, a consistent estimator for Φ is given by

$$\frac{1}{n}\sum_{i=1}^{n}\left[X_i - E_n\left(X|Z_i\right)\right]\left[X_i - E_n\left(X|Z_i\right)\right]' \mathbb{I}\left(|f_n(Z_i)| > b\right) ,$$

where $f_n(\cdot)$ is the kernel estimator of the density of Z and b is a trimming parameter, that is necessary in theory but does not play a crucial role in practice in our case.

As suggested by Robinson (1988), under conditional heteroscedasticity of unknown form, $\widehat{\lambda}$ will still be $\sqrt{n}$-consistent and, following Eicker (1963), a consistent estimator of its asymptotic variance-covariance matrix can be obtained using the estimated residuals $\widehat{U}_i$ as

$$\frac{1}{n}\sum_{i=1}^{n}\left[X_i - E_n\left(X|Z_i\right)\right]\left[X_i - E_n\left(X|Z_i\right)\right]' \widehat{U}_i^2 \mathbb{I}(|f_n(Z_i)| > b)\,.$$

Appendix C. Specification checks by kernel methods

Consider the model (11) and estimated residuals $\widehat{U}_i$. Fan and Li (1996) propose to check the specification of the model by considering the statistic

$$V_n = \frac{2}{n(n-1)}\sum_{i<j}\widehat{U}_i\widehat{U}_j\frac{1}{h^p}K\left(\frac{X_i - X_j}{h}, \frac{Z_i - Z_j}{h}\right)\,,$$

where h is a bandwidth, p is the number of components in (X, Z) and $K(\cdot,\cdot)$ is a kernel from $\mathbb{R}^p$ to $\mathbb{R}$. The statistic is such that, under $H_0 : E[U|X, Z] = 0$,

$$nh^{p/2}V_n \xrightarrow{d} N(0,\omega^2), \quad \text{with } \omega^2 = 2E\left[E^2(U^2|X,Z)f(X,Z)\right]\int K^2(u)\,du\,.$$

Under any alternative to the null, i.e. under any misspecification of the regression model, $nh^{p/2}V_n \xrightarrow{p} +\infty$. This means that asymptotically, V_n can take negative values only under the null hypothesis. Hence, a one-sided normal test can be based on $nh^{p/2}V_n/\omega_n$, where ω_n^2 is a consistent estimator of ω^2 as

$$\omega_n^2 = \frac{4}{n(n-1)}\sum_{i<j}\widehat{U}_i^2\widehat{U}_j^2\frac{1}{h^p}K^2\left(\frac{X_i - X_j}{h}, \frac{Z_i - Z_j}{h}\right)\,.$$

Suppose now we would like to test the specification of (11) with respect to X only, that is $H_0 : E[U|X] = 0$. Then we can consider the statistic

$$V_n^* = \frac{2}{n(n-1)}\sum_{i<j}\widehat{U}_i\widehat{U}_j\frac{1}{h^q}L\left(\frac{X_i - X_j}{h}\right),$$

where h is a bandwidth, q is the dimension of X and $L(\cdot)$ is a kernel from $\mathbb{R}^q$ to $\mathbb{R}$. A generalization of results of either Fan and Li (1996) or Delgado, Dominguez and Lavergne's (2000) show that this statistic has a behavior similar to V_n, i.e. under $H_0 : E[U|X] = 0$,

$$nh^{p/2}V_n^* \xrightarrow{d} N(0,\omega^{2*}), \quad \text{with } \omega^{2*} = 2E\left[E^2(U^2|X)f(X)\right]\int K^2(u)\,du,$$

and $nh^{p/2}V_n^* \xrightarrow{p} +\infty$ under any alternative. As before, a one-sided normal test relies upon $nh^{p/2}V_n^*/\omega_n^*$, where ω_n^{*2} is a consistent estimator of ω^{*2} similar to ω_n^2.

Table 1: Descriptive statistics (320 observations)

Variable	Mean	St. deviation	Minimum	Maximum
δ	0.5032	0.2875	0.0024	0.9960
t	223.82	68.68	91.01	561.06
B	359.74	515.23	4.00	5793
$PART$	0.4278	0.3128	0.0037	0.9982

B: gross Biological Oxygen Demand (BOD) emission level (in kg. per day); δ: BOD abatement rate (in %); t: BOD emission charge (in French Francs); $PART$: indicator of emissions toxicity (in %).

Table 2: Semiparametric regression

	Ordinary Least Squares		Instrumental variables	
Variable	Estimate	St. dev.	Estimate	St. dev.
$\log t$	0.6830**	0.1910	0.6729**	0.1933
$\log B$	0.0050	0.0522	-0.0179	0.0612
Food and Drinks	-0.0845	0.2953	-0.0775	0.3062
Dairy and Meat Products	-0.2155	0.2856	-0.2000	0.3124
Chemicals	-0.0031	0.3153	0.0125	0.3146
Iron and Steel	-0.4101*	0.2420	-0.4207	0.3243
Paper and Wood	-0.2919	0.3201	-0.2778	0.3665
α_0	0.9927**	0.0723	1.0266**	0.0917
α_1	2.4641**	0.4151	2.4861**	0.4270

* and ** indicate significant parameters at the 10% and 1% level respectively.

Table 3: Specification tests

	Parametric part		Full model	
a	Test statistic	P-value	Test statistic	P-value
0.50	-0.9732	0.8348	-1.2312	0.8909
0.75	-1.2296	0.8906	-0.4743	0.6824
1.00	-1.6503	0.9506	-0.1774	0.5704
1.25	-1.4834	0.9310	0.3195	0.3747
1.50	-1.3830	0.9167	0.3378	0.3678

a is a constant allowing the bandwidth parameters to vary.

Figure 1: Status-Quo δ^0, First-Best δ^* and Second-Best δ^{**} abatement rates.

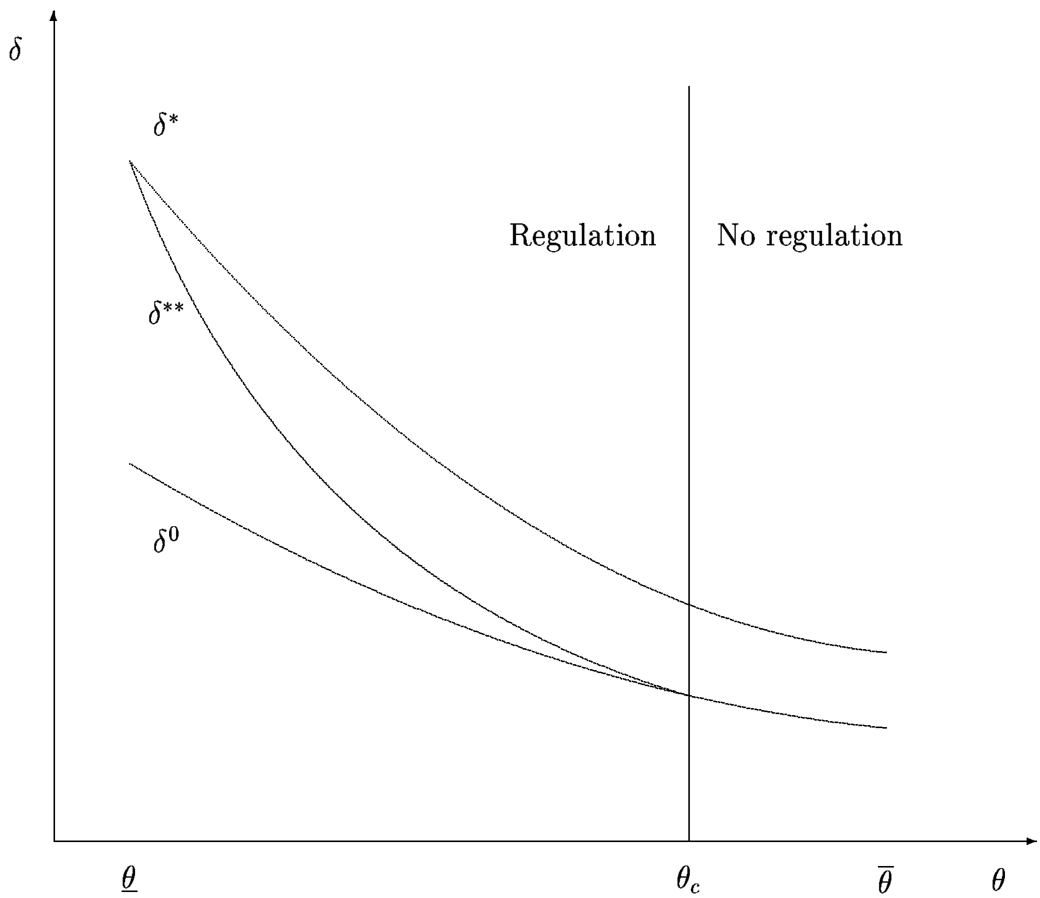

Figure 2: Function g(.)