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Regulatory Constraints and Cost Efficiency of the Italian Public Transit Systems: An Exploratory Stochastic Frontier Model ♣

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Abstract

The core of the present study is to analyse the cost efficiency of public transit systems in Italy, investigating the way subsidization mechanisms and network characteristics affect the x-inefficiency levels of firms. The analysis uses a seven-year (1993-1999) unbalanced panel of 45 Italian public-owned companies run under two different regulatory schemes (cost-plus or fixed-price). A stochastic cost frontier is estimated by applying the Battese and Coelli (1995) econometric model, which allows the inefficiency measures to depend on observable factors that are firm-specific and can vary over time. The results provide some insights regarding the appropriate mechanism of granting subsidies, and in turn they are useful for assessing the ongoing reform of the sector. The evidence indicates that there is scope for transport policy to increase cost efficiency. Efforts have to be intensified in the twofold direction of replacing cost-plus subsidization mechanisms with high-powered incentive schemes and improving exogenous operating conditions of the network. The latter goal could be pursued by acting on factors such as the road space allocation (exclusive lanes for trams and buses, parking and private traffic regulation), or the provision of incentives for the use of public modes.

Key words: Local public transport, Cost efficiency, Regulation, Stochastic frontier

JEL: C13, C24, L51, L92, R41

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1. Introduction

During the first half of the nineties, many efforts were made in the attempt to redress the ruined financial and economic situation characterizing the Italian local public transit (LPT) industry. Nevertheless, these interventions were only stopgap measures, that turned out to be inadequate to achieve the general goal of a structural readjustment of accounts. A legislative reform started with Law 549/1995, which introduced the financial responsibility for the Regions, and subsequently continued with the *Decreti Legislativi* 422/1997 and 400/1999. These normative actions met the need of a deep shake-up of the entire industry, that many experts indicated as the only way to achieve a remarkable improvement in terms of productive efficiency and effectiveness of the service.

The core of the present study is to put forward information on x-efficiency (Leibenstein, 1966) of the Italian public transit systems, so as to highlight distortions from the best-practice behavior of cost minimization. Moreover, in view of the importance of regulatory constraints in the production analysis of public utilities, as emphasized by recent empirical literature[1], we investigate how subsidization mechanisms affect firms' efficiency levels. The results of this analysis, which includes companies operating under two different regulatory schemes (cost-plus or fixed-price), should provide some insights regarding the appropriate mechanism of granting subsidies, and in turn they might be useful to assess the ongoing reform of the sector.

A seven-year unbalanced panel data of 45 Italian public-owned LPT companies is used in the empirical analysis. In the light of purposes of this work, the observed time period (1993-1999) is particularly informative, since it encompasses both years before and after the start-up of the reform. The estimation of a stochastic cost frontier model is carried out by applying the Battese and Coelli (1995) methodology, which assumes the inefficiency terms to be a function of a set of explanatory variables including firm-specific and time effects. In particular, the present investigation analyses how regulation, network characteristics, and their interaction affect cost inefficiency.

The rest of the paper is organized as follows. Section 2 briefly summarizes the regulation of the Italian LPT industry in the last decade, focusing on the subsidization schemes and the related incentive mechanisms. In Section 3, we develop the econometric model. Section 3.1, in particular, specifies the stochastic cost frontier,

[1] For the LPT industry, see Dalen & Gomez-Lobo (1997), Ivaldi (1997) and Gagnepain & Ivaldi (1998).

while Section 3.2 deals with the modeling of x-inefficiency effects. The database is described in Section 4. Section 5 comments on the empirical results, discussing both the technology properties (Section 5.1) and the evidence on cost inefficiency (Section 5.2). Section 6 summarizes the major findings and provides some policy indications.

2. Subsidization mechanisms and incentives

A common feature of the regulatory framework of public transit systems in most countries is the provision for transfers from the local authority to the LPT firm. Since the latter face universal service obligations, commercial revenues are generally not high enough to cover operating costs. The payment of a subsidy is then required to ensure the balance of the budget. In 1995 the share of public subsidies over operating costs for the Italian bus-line companies amounted to about 71%. The LPT industry in Italy has been interested by several important regulatory interventions during the last seven years, in the effort to reduce the waste of public funds spent on collective transport.[2] In 1995, Law 549 implied the abolition of the old system of redressing deficits of LPT firms through resources drawn on the National Transport Fund, a central government grant system properly created for this purpose. The opportunity cost of public funds was thereby transferred to the Regions, who are nowadays in charge of the programming of services. Subsequently the Reform has been implemented by the *Decreti Legislativi* 422/1997 and 400/1999.

An important innovation that the legislator tried to introduce in the organization of local public transport is the increase of the financial responsibility of all the subjects operating in the sector, i.e., local authorities and LPT firms. The purpose is to better select which public service deserves to receive subsidies[3] and to stimulate the recovery of productive efficiency by transportation companies. Here the necessity comes to eliminate the transfers from the central government and to replace them with forms of taxation at regional level, in order to make binding for local authorities any measure for an efficient use of public resources. In parallel, the reform dictates that the relations between the regulatory subject and the transit service provider are governed through the so-called *service contract*, a formal agreement which defines the rules that the LPT

2 The Italian regulatory framework is analyzed in detail in Piacenza (2000b) and Boitani and Cambini (2001a).

3 *Decreto Legislativo* 422/1997 names these categories “minimum services”. In practice, the definition of minimum service should correspond to the level of service that a community wants to make universally and actually affordable to each of its member, normally at non-market special tariff conditions.

company must obey as well as the reimbursement and risk sharing scheme between the regulator and the operator[4].

After the enactment of Law 549, subsidization practices began to develop differently in Italy. Before 1996, all LPT systems were run under cost-plus regimes, characterized by the full recovery of budget losses by local authorities[5]. According to this scheme, known in the regulatory practice as *management contract* (European Commission, 1998), the operator does not bear any risks on costs (industrial risk) and revenue (commercial risk). Thus, in the light of the new theory of regulation (Laffont and Tirole, 1993), it has no incentives to produce efficiently. From 1996 onwards, some municipalities introduced alternative reimbursement mechanisms that, even if not formalized within a proper service contract yet, virtually overcame the ex-post balancing of accounts: the *gross cost schemes*, under which the industrial risk is entirely borne by the operator while the commercial risk is borne by local authority, and the *net cost schemes*, that provide for the assumption of both types of risk by the company. These two types of contractual arrangements are traceable to what the theory of incentives in regulation names fixed-price schemes. In both cases, the transfer from the local authority is defined ex-ante, on the basis of expected operating costs (gross cost approach) or expected operating deficits (net cost approach), and realized costs/deficits that deviate from the fixed criteria will not influence the level of subsidies. Thus, compared to companies under the traditional cost-plus regime, the operators facing fixed-price mechanisms are assumed to confront high-powered incentives towards a cost minimizing behavior.

It is worthwhile to underline that both cost-plus and fixed-price schemes are not optimal rules in the sense specified by the new theory of regulation. According to this approach, because of the presence of informational constraints, optimal mechanisms must solve the trade-off between the efficiency incentives typical of fixed-price schemes and the rent extracting properties of a cost-plus regulation[6]. The complex problem of designing an optimal contract is out of the scope of our study[7], since only fixed-price or cost-plus schemes are practiced at the present time in the Italian LPT industry. Given the above discussion on the two regulatory mechanisms, the present paper is aimed at investigating if transit companies run under fixed-price regimes are more cost efficient due to the fact that they face stronger incentives to increase

4 The contents of the service contract are define in detail by article 19 of *Decreto Legislativo* 422/1997.

5 An exception is represented by Bolzano ACT-VVB, already subjected to a standard-cost regime from 1989.

6 See Laffont and Tirole (1993) for a complete description of this problem.

7 In this regard, see Wunsch (1994), Dalen and Gomez-Lobo (1995, 1997) and Gagnepain and Ivaldi (1999), and for an application to the Italian regulatory framework, Boitani and Cambini (2001b).

managerial effort[8]. If it is so, then we may conclude that incentive theory and modern regulatory economics are necessary components in the production analysis of regulated utilities. On the policy side, this investigation allows to assess if the subsidization schemes recently introduced in Italy are suitable in order to recover efficiency, which is one of the goals pursued by the legislative reform.

3. The econometric model

The frontier concept arises in the econometric practice when one considers that theoretical production and cost functions represent the maximum and minimum values, respectively, of an optimization problem. In this sense the notion of cost function may be interpreted as a frontier relationship, i.e., a benchmark behavior, because it is impossible for a firm to achieve costs lower than the minimum requirement, whereas higher levels are often observed in the real world, which reveal the presence of *x-inefficiency* in the production process.

In the case of unbalanced panel data the general stochastic frontier specification for a *variable cost function* can be written as:

$$VC_{ft} = VC(Y_{ft}, P_{ft}, Z_{ft}, \tau_{ft}; \beta)\exp\{\psi_{ft}\}, \quad [1]$$

with $\psi_{ft} = v_{ft} + u_{ft}$, [2]

where VC denotes variable cost, Y represents a vector of output, P is an $m \times 1$ vector of prices of variable factors, Z is an $n \times 1$ vector of variables including quasi-fixed inputs and network characteristics, τ indicates the year of the observation involved, and β is a $k \times 1$ vector of technology parameters to be estimated. For all variables the subscript f indexes firm ($f = 1, \ldots, F$), and t indexes observation ($t = 1, \ldots, T_f$). The f subscript on T is used to indicate the unbalanced nature of the panel. For all f $1 \leq T_f \leq T$, with T denoting the maximum number of observations available for a firm[9].

As usually in the frontier literature, the error term ψ_{ft} is decomposed into two components: (*i*) the white noise component, v_{ft}, which capture the effects of all exogenous shocks to the production process and (*ii*) the inefficiency term, u_{ft}, representing firm- and time-specific cost inefficiency. The statistical noise term, v_{ft}, makes the frontier cost function $VC(.)$ stochastic and can take both positive and negative

[8] A preliminary analysis in this direction, based on average cost and productivity indicators, is provided in Fraquelli et al. (2001a).

[9] Although it is assumed that there are T time periods for which observations are available for at least one of the F firms involved, it is not necessary that all the firms are observed for all T periods in an unbalanced panel data specification of the econometric model.

values, according to whether the exogenous shocks have unfavorable or favorable effects on cost. The *non-negative* error component, u_{ft}, on the other hand, indicates the amount by which the logarithm of cost of the f^{th} firm at the t^{th} observation exceeds the logarithm of stochastic frontier, $\ln VC(.) + v_{ft}$, due to x-inefficiency. When $u_{ft} = 0$ for a particular firm, *f*, at observation *t*, it attains the cost frontier.

3.1. Specification of the stochastic frontier cost function

To analyze the productive structure of the Italian LPT industry we chose a variable operating cost model. The fixed assets investments in this sector are strictly related to government financing programs, so it is not proper to suppose that companies exhibit a cost-minimizing behavior with respect to capital too. Therefore, as Windle (1988), Levaggi (1994) and Fabbri (1998) suggest, the rolling stock should be considered as a fixed factor in the short-run. The model includes: a scalar output (Y); the prices of three variable factors, i.e. labor (L), fuel (F), materials and services (MS); a quasi-fixed input (K); a network characteristics (SP), i.e. the average commercial speed; a time trend variable (τ).

We use a composite measure of the output to reflect the global productive structure of firms. It is well-known in transportation literature that the output definition is a much debated question, since it can lead to different results, for example in terms of scale economies. The output indicator is computed by multiplying the transit firm's fleet size, measured in terms of total places offered[10], and the total traveled kilometers. We want to point out some remarks about this kind of output. If we consider the operative context of the LPT industry, a firm must supply the service on a certain number of lines, offering a certain number of places and trips on this network. Our definition of output allows us to take into account the length of the network, the frequency of the service and the size of the fleet. Furthermore, this measure is particularly suitable to our specific firm sample, which includes both urban and extra-urban services. As it was not possible to separate the urban activity from the extra-urban one, we defined an aggregate output and aimed to weight their specific characteristics[11].

[10] The total places offered were calculated by multiplying the number of vehicles owned by each unit and their average load capacity.

[11] Generally speaking, the extra-urban firms can perform a higher number of kilometers than the urban units, covering a larger network, but the operative context is very different (a lower number of passengers, longer trips, different traffic conditions). On the other hand, a urban company reasonably offers a higher number of places (buses are larger and also their number is higher, because there is a more intensive demand to satisfy).

The capital stock plays the role of quasi-fixed input in our variable cost model. It is represented by the number of vehicles owned by LPT companies weighted by the average fleet age. We calculated the indicator as follows: K_{ft} = (*number of vehicles in the rolling stock*)*(age_s / age_{ft}), where age_s is the average fleet age in the whole analyzed sample, while age_{ft} is the average fleet age of firm *f* at the observation *t*.

Prices of variable factors were computed getting information from the balance-sheet statistics. The labor price (P_L) was obtained by dividing total labor costs by the average annual number of service workers (drivers, maintenance workers and administrative staff). The average price of fuel (P_F) was obtained by dividing fuel costs by the annual number of liters consumed[12]. Expenses for materials and services represent a residual cost category. It has been divided by the annual seat-kilometers offered to obtain an average price for this input (P_{MS})[13]. Indeed, it is reasonable to assume that this kind of expense strictly depends on the actual exploitation of the network.

In addition to the standard variables of a proper cost function, we included in the model the average commercial speed of LPT vehicles (*SP*), already considered in some works on the industry (for instance, Windle, 1988; Levaggi, 1994; Wunsch, 1996; Gagnepain, 1998). The specificity of the territorial area where the service is provided makes it difficult to compare the cost performance of different firms. Indeed, the traffic conditions and the geographical characteristics are peculiar to each network. To some extent, the average commercial speed should reflect differences in these environmental factors. Incorporating the variable into the cost frontier, costs are expected to lower with increasing network speed.

We added to the model a time trend too, measured in years, so as to account for possible effects of Hicks neutral technological change. In fact, given the seven-year length of the panel the impact of scientific or organizational progress should not be negligible. Other things unchanged, costs are then expected to diminish over time.

A translog functional form is chosen for this analysis[14]. The stochastic frontier cost model [1]-[2] is then defined by equation [3]:

12 For a few firms which utilize tramways, trolley-lines or railways and consume electricity, kilowatt-hours were transformed in equivalent-liters.

13 Seat-kilometers are the multiplication of traveled kilometers by the average load capacity of vehicles.

14 Given the regularity conditions ensuring duality, the estimation of a translog cost function does not impose any other a priori restriction on the characteristics of the below technology. In particular, the elasticity of substitution and the returns to scale can vary with both the output level and the combination of inputs. This fully satisfies the criterion of model generality.

$$\ln\left(\frac{VC_{ft}}{P_{Fft}}\right)=\beta_0+\beta_y\ln Y_{ft}+\beta_k\ln K_{ft}+\sum_i\beta_i\ln\left(\frac{P_{i\,ft}}{P_{Fft}}\right)+\beta_{SP}\ln SP_{ft}+\sum_i\beta_{iy}\ln\left(\frac{P_{i\,ft}}{P_{Fft}}\right)\ln Y_{ft}+$$
$$+\sum_i\beta_{ik}\ln\left(\frac{P_{i\,ft}}{P_{Fft}}\right)\ln K+\sum_i\beta_{iSP}\ln\left(\frac{P_{i\,ft}}{P_{Fft}}\right)\ln SP_{ft}+\beta_{yk}\ln Y_{ft}*\ln K_{ft}+$$
$$+\beta_{ySP}\ln Y_{ft}*\ln SP_{ft}+\beta_{kSP}\ln K_{ft}*\ln SP_{ft}+\frac{1}{2}\beta_{yy}(\ln Y_{ft})^2+\frac{1}{2}\beta_{kk}(\ln K_{ft})^2+$$
$$+\frac{1}{2}\beta_{SPSP}(\ln SP_{ft})^2+\frac{1}{2}\sum_i\sum_j\beta_{ij}\ln\left(\frac{P_{i\,ft}}{P_{Fft}}\right)\ln\left(\frac{P_{j\,ft}}{P_{Fft}}\right)+\beta_\tau\tau_{ft}+v_{ft}+u_{ft}$$
$$i,j\in\{L,MS\}, \qquad [3]$$

where the normalization of the monetary variables, *VC*, P_L and P_{MS}, with respect to the price of fuel, P_F, is made to ensure the linear homogeneity of the cost function in input prices[15].

The x-inefficiency term, u_{ft}, reflects the inability of firm *f* at the observation *t* to attain the potential minimum cost defined by the stochastic frontier. The specification for this effect and the discussion of the estimation technique for the final stochastic frontier model are given in the next two sections.

3.2. Modeling inefficiency effects

Several innovations concerning the estimation of inefficiency using the stochastic production and cost frontier approach have been introduced since the pioneer contributions of Aigner, Lovell and Schmidt (1977) and Meeusen and van den Broeck (1977)[16].

Researchers have attempted to overcome the shortcomings present in the above frontier models: by specifying distributional forms for the inefficiency effects more general than the half-normal and exponential distributions (Stevenson, 1980; Greene, 1990)[17]; proposing functional forms alternative to the traditional Cobb-Douglas

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15 Symmetry property (β_{ij} = β_{ji} for all *i*, *j*) is also imposed a priori, whereas the other regularity conditions, viz., monotonicity of the cost function in input prices and output, and concavity in input prices are checked ex-post.

16 A brief introduction to the literature on stochastic frontier modeling and efficiency measurement is provided in Piacenza (2000a). For a recent and more detailed review see Kumbhakar and Lovell (2000).

17 A common criticism of the stochastic frontier method is that there is no a priori justification for the selection of any particular distributional form for the inefficiency effects, u_{ft}. The half-normal and the exponential distributions are arbitrary selections. Since both of these distributions have a mode at zero, it implies that there is the highest probability that the inefficiency effects are in the neighborhood of zero. This, in turn, implies relatively high efficiency. In practice, it may be possible to have a few very efficient firms, but a lot of quite inefficient firms.

technology (e.g., Greene, 1980b; Kumbhakar, Ghosh and McGuckin, 1991)[18]; extending the analysis to the dual cost function (e.g., Schmidt and Lovell, 1979; Ferrier and Lovell, 1990; Kumbhakar, 1991)[19]; accommodating panel data (e.g., Battese and Coelli, 1988; Cornwell, Schmidt and Sickles, 1990; Kumbhakar, 1990; Battese and Coelli, 1992; Lee and Schmidt, 1993)[20].

More importantly for the purpose of this work, a number of later empirical studies (e.g., Pitt and lee, 1981; Kalirajan, 1981; Kalirajan and Shand, 1989; Mester, 1997) have investigated the *determinants* of productive inefficiencies among firms in an industry by regressing the predicted inefficiency effects, obtained from an estimated stochastic frontier, upon a vector of firm-specific factors, such as the degree of competitive pressure, input and output quality indicators, various managerial characteristics, etc., in a second-stage analysis. There is, however, a significant problem with this two-stage approach. In the first stage, the inefficiency effects are assumed to be independently and *identically* distributed in order to use the approach of Jondrow et al. (1982) to predict the values of the technical inefficiency effects. However, in the second stage, the predicted inefficiency effects are assumed to be a function of a number of firm-specific factors, which implies that they are *not* identically distributed, unless all the coefficients of the factors are simultaneously equal to zero.

Kumbhakar, Ghosh and McGuckin (1991) and Reifschneider and Stevenson (1991) noted the above inconsistency and specified stochastic frontier models in which the inefficiency effects were defined as explicit functions of some firm-specific factors, and all parameters were estimated in a single-stage maximum likelihood (ML) procedure. Huang and Liu (1994) also presented a model for a stochastic frontier

[18] The Cobb-Douglas functional form has been commonly used in the empirical estimation of frontier models. Its simplicity is a very attractive feature. This simplicity, however, is associated with a number of restrictive properties. The Cobb-Douglas technology exhibit the same value of returns to scale for all firms in the sample. Further, the elasticities of substitution between productive factors are equal to one.

[19] The cost frontier approach appears to be a significant improvement to the efficiency analysis. It accounts for the possibility of exogenous output and endogenous inputs, permits the measurement of technical and allocative inefficiency, and can be easily extended to account for multiple outputs. Further, the objective of (total or variable) cost minimisation may often be a proper assumption. It is particularly suitable in environments where output is demand driven, and so also can be considered to be exogenous. Many regulated industries, such as electricity generation, gas distribution, or public transit service, satisfy these exogeneity criteria. Moreover, in many industries output is not storable, and so the output maximization objective that underlies the estimation of output-oriented technical efficiency would be inappropriate.

[20] Panel data have some advantages over cross-sectional data in the estimation of stochastic frontier models. First, when panel data are available there is no need to specify a particular distribution for the inefficiency effects, because the parameters of the model can be estimated using the traditional panel data techniques of fixed-effects (dummy variables) or random-effects. Second, also by proceeding with the more commonly used maximum likelihood (ML) methods, the availability of panel data generally implies that there are a larger number of degrees of freedom for the estimation of parameters. Third, panel data permit the investigation of efficiency change over time.

production function, in which the non-negative technical inefficiency effects were a linear function of variables involving firm characteristics, together with their interactions with the input variables of the frontier function[21]. Battese and Coelli (1993, 1995) extended these approaches model to accommodate panel data, which permits to include both firm-specific and time effects in the model adopted to explain inefficiencies.

Coelli (1996) wrote a computer program, FRONTIER Version 4.1, that automates the ML method for estimation of the parameters of Battese and Coelli (1995) model and also allows to specify the stochastic frontier in terms of a dual cost relationship instead of a production function. The availability of such a program, which permits to easily estimate a stochastic cost frontier and analyze the sources of x-inefficiency, further allowing unbalanced panel data (as the sample we have at hand), induced us to apply the Battese and Coelli (1995) model for studying the effects of regulatory constraints on the cost efficiency of public transit systems in Italy[22]. The objective, in particular, is to investigate whether the predictions from the theory of incentives in regulation (Laffont and Tirole, 1993) help explain differences in productive efficiency, i.e. *do high powered incentive regulatory schemes increase efficiency as compared to low powered schemes?* To do this, we start from issues of a stochastic cost frontier model recently developed by Gagnepain and Ivaldi (1998)[23].

[21] This makes their model a non-neutral shift of the traditional average response function, in that the marginal products of inputs and marginal rates of technical substitution (MRTS) depend on the firm-specific variables in the inefficiency model.

[22] A ML systems estimator, involving the cost frontier and the factor-share equations, would provide more efficient estimators of the parameters of a cost function than the single-equation estimator automated in FRONTIER. This system approach also has the advantage of explicitly accounting for allocative inefficiency (reflected in the error terms of the factor-share equations), that represent violations of the first-order conditions for cost minimization. However, a frontier systems estimator suffers from some problems. First, it is not as yet automated in any computer package, hence one would need to write code for it in some way (using SAS, GAUSS or TSP) that is often very time-consuming. Second, once one specifies flexible functional forms, such as the translog form, where the implied production function cannot be derived, the decomposition of the overall cost efficiency into technical and allocative components (what has come to be known as “the Greene problem”) requires some restrictive assumptions, and none of the existing approaches (e.g., Ferrier and Lovell, 1990; Kumbhakar, 1991; Mensah, 1994; Kumbhakar, 1997) is exempt from criticism from some quarter. Further, estimation problems often arise when one tries to numerically solve the rather complicated likelihood functions that are involved.

[23] These authors directly incorporate into the cost minimization problem the distortions on productive activity due to regulatory constraints and the presence of informational asymmetries in the regulator-firm interaction. In this way, the error component the literature generally attributes to cost inefficiency, u_{ft}, is already built-in and the econometric frontier model exactly coincides with the theoretical cost model, without the necessity of adding other more than a random disturbance term to account for exogenous shocks and potential measurement faults. In this work we do not utilize the economic-theory-based methodology employed by Gagnepain and Ivaldi (1998). However, the present study constitutes a useful exploratory analysis that could provide insights about the impact on cost efficiency of the investigated variables. Thus, it represents the starting base for future more elaborated works.

We postpone to the following section the discussion concerning statistical properties of the Battese and Coelli (1995) model and the estimation procedure. In succession, the principal features of the Gagnepain and Ivaldi (1998) model are briefly resumed, focusing the attention on the role of *informational asymmetries* and incentives in the regulator-firm relationship. We describe then the variables we suppose to capture these effects within the methodological framework proposed by Battese and Coelli.

From equations [1]-[2] *overall cost inefficiency* could be defined by the expression $\exp\{u_{ft}\}$, which represents the extent to which observed cost, VC_{ft}, exceeds the frontier minimum level after accounting for the effect of statistical noise, $VC(Y_{ft}, P_{ft}, Z_{ft}, \tau_{ft}; \beta)\exp\{v_{ft}\}$. Gagnepain and Ivaldi specified u_{ft} as $g(\theta_{ft}, e_{ft})$, so that $\exp\{u_{ft}\} = \exp\{g(\theta_{ft}, e_{ft})\}$, where g is a function strictly increasing in θ and decreasing in e[24], θ is a parameter reflecting the presence of exogenous *technical inefficiency*, source of a fall in the productivity of labor input[25], and e, that could be thought as more responsible for *allocative inefficiency*[26], represents the cost reducing activity exerted by managers to counterbalance the effect of the intrinsic lack of labor productivity[27].

The intrinsic inefficiency level, θ, is supposed to be perfectly known by the firm and not known or imperfectly known by the regulatory authority (*adverse selection* phenomenon), as the latter does not take part in the production process. On the other hand, it should be even harder for the regulator-principal to observe and directly control the effort provided by the manager-agent, e (*moral hazard* phenomenon)[28]. Since there are no clear evidences on the motivation of a manager to work as hard as he could, it is where the type of regulatory scheme and related incentives faced by the firm during the production process play a role in reducing overall cost inefficiency, $\exp\{g(\theta, e)\}$.

The authors elaborated a structural cost frontier model, in which the cost reducing effort of the producer is endogenous in the sense that it depends on the regulatory environment impinging on its activity. By applying their model to the analysis of public transit systems in France (Gagnepain and Ivaldi, 1998 and 1999) they showed that operators subjected to fixed-price subsidization schemes exhibited higher effort levels compared to ones displayed by companies run under cost-plus mechanisms[29]. Given the

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[24] Subscripts, f and t, are omitted for convenience in the presentation.

[25] This type of inefficiency is given and cannot be changed by the management in the short run. It depends on factors such as geographical and historical characteristics of a network, the structure of the labor force or the ability level of drivers.

[26] That is, the failure (attributable to an insufficient managerial effort) in making the actual MRTS between any two inputs equal to the corresponding input price ratio.

[27] The operator may spend time and effort in monitoring workers, for instance providing drivers with training programs, solving potential conflicts among them, avoiding strikes, etc.

[28] Indeed, the regulator cannot distinguish between the effect of intrinsic inefficiency and the impact of cost reducing effort.

[29] This evidence is clearly consistent with the new theory of regulation (Laffont and Tirole, 1993) which

level of intrinsic inefficiency, this implied lower cost distortions over the best-practice frontier, and the fact that fixed-price contracts provide more incentive for efficiency has been proved. Only for a group of operators characterized by a fairly high technical inefficiency, contractual arrangements did not appear to be very relevant for the firm's productive performance. Indeed, in these cases the cost reducing activity exerted by managers had a little weight in determining global cost inefficiency. Consequently, the distortions over the frontier remained significant also in presence of high powered incentive schemes.

Taking cue from the above results, the Battese and Coelli (1995) stochastic frontier model is applied to the analysis of overall cost inefficiency, or x-inefficiency, of public transit systems in Italy. The emphasis is put on the role played by the regulatory context, in terms of the different subsidization mechanisms that companies has to face, taking also into account the specific characteristics of each network in a way which will be specified later.

Battese and Coelli define the inefficiency effects, u_{ft}, as non-negative random variables assumed to be a function of a set of firm-specific explanatory variables which may vary over time, z_{ft}, and an unknown vector of coefficients, δ, associated with the $z_{ft}s$. The explanatory variables in the inefficiency model would be expected to include any factors that help explain the extent to which the variable cost observations exceed the corresponding stochastic frontier cost values, $VC(Y_{ft}, P_{ft}, Z_{ft}, \tau_{ft}; \beta)\exp\{v_{ft}\}$. The z_{ft}-vectors usually have the first element equal to one[30] and may also include some variables involved in the cost function[31] and/or interactions between these latter and firm-specific factors.

The x-inefficiency effect, u_{ft}, incorporated in the composed error term, ψ_{ft}, of the general stochastic frontier model [1]-[2] could be specified by equation [4],

$$u_{ft} = \delta' z_{ft} + w_{ft} = \sum_q \delta_q z_{qft} + w_{ft} \,, \qquad [4]$$

where the q subscript on δ and z_{ft} indexes explanatory variables ($q = 0, \ldots, Q$), and w_{ft} is a random variable making the inefficiency effect stochastic whose distribution will be defined in the next section.

defines the fixed-price schemes as the maximal powered incentive contracts. More precisely, the optimal effort level equalizes the marginal disutility of effort and the marginal cost savings under fixed-price regimes while it is nil under cost-plus regimes.

30 Not including an intercept parameter, δ_0, in the z_{ft}-vectors may result in the estimators for the δ-parameters being biased and the shape of the distributions of the inefficiency effects, u_{ft}, being unnecessarily restricted.

31 Provided the inefficiency effects are stochastic.

Since our objective is to verify if the causes of cost inefficiency affecting the Italian public transit systems should be searched in the system of incentives generated by the regulatory environment, we first introduce a regulation dummy, R_{ft}, as determinant of u_{ft}.

As mentioned before two great categories of reimbursement rules are observed in practice: cost-plus schemes, according to which subsidies are paid by the local authority to the company so to allow ex-post budgets to be balanced[32], and fixed-price schemes, where the transit operator obtains a transfer defined ex-ante in order to finance an expected operating deficit[33]. Variable R_{ft} takes value 0 when cost-plus regulation is observed, and value 1 in cases where fixed-price schemes are applied. According to the predictions from the theory of incentives in regulation and the evidence emerged in the study of Gagnepain and Ivaldi, cost distortions over the frontier, $\exp\{u_{ft}\}$, are expected to be significantly lower under fixed-price regulation, as in such a context the company's manager should increase the effort to reduce production costs. The sign of the parameter associated with the regulation dummy, δ_R, is then expected to be negative.

The second important explanatory variable we include in the specification of the cost inefficiency model [4] aims at capturing the effects on x-efficiency attributable to the specific operating conditions of the environment where the transit service is provided. We refer to factors as the geographical and historical characteristics influencing the structure and the operability of the network, the ability level of drivers, the public policy for local mobility, etc. These are the elements that in the Gagnepain and Ivaldi model contribute to determine the intrinsic labor inefficiency level, θ_{ft}, what the authors call "exogenous technical inefficiency".

Here we do not obtain a specific estimate for this unobservable component of the global cost inefficiency[34]. Still an attempt is made to include in the inefficiency model [4] a variable strictly related to the above factors (from now on "network characteristics"), likely to influence the exogenous technical inefficiency and then the level of overall cost distortion. To this end we introduce as a proxy for network

32 In Section 2 we refer to this reimbursement rule as the "management contract", under which the regulated firm does not bear any risk.

33 This type of reimbursement rule specifically refers to the subsidization scheme we have called before "net cost contract". Actually, we have seen that the class of fixed-price schemes also includes the "gross cost contract". Under this variant, the authority receives the commercial revenue and pays the firm's expected costs. In terms of incentives to produce efficiently, it is similar to the first variant of fixed-price contracts. Under the "net cost contract" option, however, the LPT operator bear all the risks on costs (industrial risk) and revenue (commercial risk), whereas under the "gross cost contract" option only the industrial risk is borne by the transit firm.

34 Indeed, we remark that θ represents an adverse selection parameter which reflects a private information on the firm's technology not known (or imperfectly known) by the regulatory authority and the econometricians.

characteristics the average commercial speed, ln SP_{ft}, already included in the specification of the frontier cost function as a network variable affecting the underlying technology[35]. Since a higher value for this variable is supposed to reflect better operating conditions[36], thus reducing the intrinsic inefficiency level, once the positive relationship between θ_{ft} and u_{ft} conjectured in the Gagnepain and Ivaldi model is assumed to be valid, we expect to find a negative sign for the coefficient associated with ln SP_{ft}, δ_{SP}, in model [4].

Furthermore, to take into account the evidence that when the intrinsic inefficiency of a network is too high, the effect of contractual arrangements on the overall cost inefficiency becomes modest[37], the impact exerted by regulatory schemes is allowed to vary with the level of average commercial speed. This is made by introducing in the model an interaction of the regulation dummy with the variable ln SP_{ft}, denoted ($R_{ft} \times$ln SP_{ft}). The relative parameter, δ_{RSP}, is expected to have a negative sign, to indicate a stronger power of fixed-price schemes in reducing x-inefficiency when regulated firms are facing more favourable exogenous operating conditions. At the same time, the negative sign of δ_{RSP} would means that the impact on cost efficiency due to a gain in the average commercial speed is strengthened in presence of fixed-price schemes, because of the higher cost reducing effort provided by managers under this type of regulation.

As the Battese e Coelli model enables us to include both firm-specific and time effects in the specification of the inefficiency model, we also incorporate in the equation [4] a time variable, τ_{ft}, indicating the year of the observation involved. It specifies that x-inefficiency may change linearly with respect to time according to the sign of the associated parameter, δ_{τ}. Given the frequent government stopgap measures adopted in the first half of the nineties to face deficits of LPT companies and the delays in bringing about the reform started with the Law n. 549 in 1995, the sign of this coefficient is expected to be positive. Moreover, an interaction between regulation dummy and time variable is also introduced, so to allow the dynamics of cost inefficiencies throughout the analyzed period to vary with the regulatory pattern. We denote this variable with ($R_{ft} \times \tau_{ft}$), while $\delta_{R\tau}$ is the relative parameter.

Under the above specifications on the set of explanatory variables, the z_{ft} s, the cost inefficiency model [4] can be written as:

[35] Transformation in logarithms is maintained for homogeneity with the equation [3]. In both cases the logarithm specification allows to interpret the partial derivatives of the dependent variables, lnVC_{ft} and u_{ft}, computed with respect to lnSP_{ft} in terms of elasticities.

[36] In fact, it is reasonable to assume that the average commercial speed may increase for instance when transit firms face more favorable geographical conditions, skilful drivers, or public policies attentive to the local traffic regulation.

[37] See the discussion above concerning the results obtained by Gagnepain and Ivaldi (1998, 1999) in their study on the French public transit systems.

$$u_{ft} = \delta_0 + \delta_R R_{ft} + \delta_{SP} \ln SP_{ft} + \delta_\tau \tau_{ft} + \delta_{RSP}(R_{ft} \times \ln SP_{ft}) + \delta_{R\tau}(R_{ft} \times \tau_{ft}) + w_{ft} \,. \quad [5]$$

Equation [5] indicates that stochastic x-inefficiency effects are assumed to be present in the frontier cost model defined by expression [3] and be linearly related to regulatory scheme and commercial speed of the transit companies, the period of observation, and the interactions of speed and time with regulation, such that an intercept parameter, δ_0, is included.

3.3. *Distributional assumptions and estimation procedure*

The final stochastic frontier model to be estimated is specified in equation [3], where the cost inefficiency effects, u_{ft}, are defined by expression [5]. According to Battese and Coelli (1995), the following distributional assumptions are made for the two components of the global error term, ψ_{ft}:

(i) the random noises $v_{ft}s$ are assumed ~ *i.i.d.* $N(0, \sigma_v^2)$, independently distributed of the cost inefficiency effects, the $u_{ft}s$;

(ii) the u_{ft} s are non-negative random variables, which are assumed to be independently but *not identically* distributed, such that u_{ft} arises from the truncation (at zero) of the normal distribution with mean $\delta' z_{ft}$ and variance σ_u^2, $N(\delta' z_{ft}, \sigma_u^2)$. This can also be written as $u_{ft} \sim N^+(\delta' z_{ft}, \sigma_u^2)$;

(iii) given the specification in model [4] for the cost inefficiency effects, the random variable w_{ft} is defined by the truncation of the normal distribution with zero mean and variance σ_u^2, such that the (variable) truncation point is $-\delta' z_{ft}$, i.e., $w_{ft} \geq -\delta' z_{ft}$. This distributional assumption on w_{ft} is consistent with the $u_{ft}s$ being ~ $N^+(\delta' z_{ft}, \sigma_u^2)$.

A truncated-normal distribution is a two-parameter distribution, with one parameter, the mode ($\delta' z_{ft}$ in the case of the u_{ft} distribution), characterizing placement[38] and the other (here σ_u^2) characterizing spread. It generalizes the one-parameter half-normal distribution, by allowing the normal distribution, which is truncated below at zero, to have a nonzero mode. Thus the truncated-normal distribution contains an additional parameter to be estimated (its mode), and so provides a somewhat more flexible representation of the pattern of inefficiency in the data.

The logic underlying the Battese and Coelli model is to relax the constant-mode property of the truncated-normal distribution specified by Stevenson (1980), by allowing the mode to be a function of a set of explanatory variables, the $z_{ft}\,s$, that are

[38] Battese and Coelli note that, unlike the Reifschneider and Stevenson (1991) model, the mode $\delta' z_{ft}$ of the u_{ft} distribution is not required here to be non-negative for each observation, so that $w_{ft} \leq 0$ is possible in a relatively unfavorable environment (i.e., if $\delta' z_{ft} > 0$).

firm-specific and vary over time, and a common vector of parameters, δ[39]. This allows cost inefficiency, which depends on the mode of the related truncated-normal distribution, to depend on exogenous observable factors[40].

The ML method is employed for simultaneous estimation of the parameters of the stochastic frontier [3] and the model for the cost inefficiency effects [5]. Using the above distributional assumptions on v_{ft} and u_{ft}, the log-likelihood function for the (T_1 + T_2 + … + T_F) sample observations, $\ln VC = (\ln VC_{11}, \ldots, \ln VC_{1T_1}; \ln VC_{21}, \ldots, \ln VC_{2T_2}; \ldots, \ln VC_{FT_F})'$, can be written as

$$L(\beta,\delta,\sigma^2,\gamma;\ln VC) = -\frac{1}{2}\left(\sum_{f=1}^{F} T_f\right)\left\{\ln 2\pi + \ln\sigma^2\right\}$$
$$-\frac{1}{2}\sum_{f=1}^{F}\sum_{t=1}^{T_f}\left\{[\ln VC_{ft} - \ln VC(Y_{ft},P_{ft},Z_{ft},\tau_{ft};\beta) - \delta' z_{ft}]^2/\sigma^2\right\}$$
$$-\frac{1}{2}\sum_{f=1}^{F}\sum_{t=1}^{T_f}\left\{\ln\Phi[d_{ft}] - \ln\Phi[d^*_{ft}]\right\}, \qquad [6]$$

where $\Phi(\cdot)$ is the standard normal cumulative distribution function, $d_{ft} = \delta' z_{ft}/(\gamma\sigma^2)^{1/2}$, and $d^*_{ft} = [(1-\gamma)\delta' z_{ft} + \gamma(\ln VC_{ft} - \ln VC(Y_{ft},P_{ft},Z_{ft},\tau_{ft};\beta))]/[\gamma(1-\gamma)\sigma^2]^{1/2}$.

The log-likelihood function is formulated in terms of the parameterization suggested by Battese and Corra (1977) who replace σ_v^2 and σ_u^2 with $\sigma^2 \equiv (\sigma_v^2 + \sigma_u^2)$ and $\gamma \equiv \sigma_u^2/(\sigma_v^2 + \sigma_u^2)$. This is done with the calculation of the maximum likelihood in mind. The parameter γ must lie between 0 and 1 and thus this range can be searched to provide a suitable starting value for use in an iterative maximization process such as the Davidon-Fletcher-Powell (DFP) routine utilized by FRONTIER Version 4.1 computer program[41]. The γ-parameterization also has the advantage to provide an indication of the relative contributions of u_{ft} and v_{ft} to ψ_{ft}. As $\gamma \to 0$ the symmetric noise component, v_{ft},

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[39] It should be noted that if the first *z*-variable has value one, as in the equation [5], and the coefficients of all other z-variables are zero, then the model reduces to the truncated-normal specification in Stevenson (1980), with δ_0 (the only element in δ) having the same interpretation as the μ parameter in Stevenson. Furthermore, if all elements of the δ-vector were equal to zero, then the inefficiency effects are not related to the *z*-variables and so the half-normal distribution originally specified by Aigner et al. (1977) would be obtained.

[40] The model is obviously a simplification, as it does not account for possible correlation structures of the cost inefficiency effects and the random errors in the frontier (the $u_{ft}s$ and the $v_{ft}s$ are assumed to be independently distributed for all $t = 1, \ldots, T_f$, and $f = 1, \ldots, F$), nor heteroskedasticity in the $u_{ft}s$ and the $v_{ft}s$. Alternative models are required to this end (see for instance Kumbhakar et al., 1995, for a model that introduce exogenous determinants of inefficiency with the variance of u_{ft} which varies across firms an over time).

[41] On the contrary, the parameterization proposed in the original contribution of Aigner et al. (1977), i.e., $\sigma^2 \equiv (\sigma_v^2 + \sigma_u^2)$ and $\lambda \equiv \sigma_u/\sigma_v$, makes the calculation of the maximum likelihood more difficult, since the λ-parameter could be any non-negative value.

dominates the one-sided cost inefficiency term, u_{ft}, in determining variation of the global residual, ψ_{ft}. The inverse occurs as $\gamma \rightarrow 1$. In the former case we are back to a traditional average function model with no stochastic x-inefficiency, whereas in the latter case we are back to a deterministic cost frontier model with no random noise included[42].

Maximization of the above log-likelihood function gives consistent and asymptotically efficient estimators of β, δ, σ^2, and γ[43]. It can be observed that: (i) the failure to include u_{ft} in the frontier cost function [3], i.e., the estimation of a traditional average response function, leads to biased and inconsistent parameter estimates, especially if the variables specified in the cost inefficiency model [4] are not orthogonal to those on the right hand side of the frontier cost function; (ii) the distributional assumptions on the $u_{ft}s$ permit the impact of Hicks neutral technological change and time-varying behavior of the inefficiency effects to be identified, in addition to the speed effects and the intercept parameters, β_0 and δ_0, in the stochastic frontier [3] and the inefficiency model [5][44]. The derivation of the log-likelihood function [6], together with the description of the procedure implemented by FRONTIER Version 4.1 to find its maximum and calculate the asymptotic standard errors of the ML estimators, are presented in the Appendix[45].

After obtaining estimates of the parameters, we consider the estimation of u_{ft}. When the model in equation [4] is assumed, the overall cost inefficiency of production for the f^{th} firm at the t^{th} observation is defined by expression [7],

$$CI_{ft} = \exp\{u_{ft}\} = \exp\{\delta' z_{ft} + w_{ft}\}, \qquad [7]$$

which takes a value between one (when $u_{ft} = 0$) and infinity (when $u_{ft} \rightarrow \infty$). The prediction of the x-inefficiencies, $C\hat{I}_{ft}$, is based on conditional expectations which generalize the estimators in Jondrow et al. (1982) and Battese and Coelli (1988). This result is also provided in the Appendix. It is worthwhile to observe that $(\delta' z_{ft} + w_{ft}) > (\delta' z_{f't} + w_{f't})$ for $f \neq f'$ does not necessary imply that $(\delta' z_{ft'} + w_{ft'}) > (\delta' z_{f't'} + w_{f't'})$ for $t \neq t'$.

42 The term *deterministic* is used because in this type of frontier models the observed cost, VC_{ft}, is bounded below by a non-stochastic (i.e. deterministic) minimum quantity, $VC(Y_{ft}, P_{ft}, Z_{ft}, \tau_{ft}; \beta)$. The models of Aigner and Chu (1968), Afriat (1972) and Schmidt (1976) are examples of deterministic frontiers.

43 See Aigner et al. (1977), p.28.

44 Time, τ_{ft}, and average commercial speed, $\ln SP_{ft}$, are assumed to affect both the production technology and the cost efficiency and occur in the data set as both variables in the frontier cost function [3] and as z-variables in the model [5] for the x-inefficiency effects.

45 The presentation follows the Appendix of Battese and Coelli (1993), excepting a few simple sign changes to take into account that we are working with a *cost* frontier instead of a *production* frontier model.

Then it follows that the same ordering of firms in terms of cost inefficiency of production does not apply to all time periods[46].

4. Data description

The study uses a seven-year unbalanced panel data of 45 companies operating in the Italian LPT sector over the period from 1993 to 1999 and are members of Federtrasporti[47]. More precisely, the panel includes 31 units observed in the years 1993-1995 and 1999 and 45 units in the period 1996-1998, for a total of 259 observations[48].

The sample composition by type of service is the following: of the 45 analyzed firms, 18 mostly operate in the urban context, 15 provide extra-urban service for the most part, and the remaining 12 have activities in both compartments. As far as the distribution by geographical area is concerned, the sample is balanced enough: 25 operators are located in the North regions and 20 in the Center-South regions (in particular, 10 in the Center and 10 in the South). The prevalence in the sample of companies providing only bus service (37 units) compared to the multi-modal firms (including also tramways, trolley-lines, and railways) reflects the modality composition at national level, where the road mode of transportation represents about 80% of LPT services in terms of seat-kilometers[49]. From the point of view of the firm size, measured in terms of average number of employed workers, the sample includes 12 firms of large size (more than 550 workers), 23 medium-sized units (151-550 workers), and 10 small operators (less than 150 workers). Finally, as far as the subsidization mechanisms are concerned, twenty-seven percent of observations (71 cases) relate to fixed-price regulatory schemes, while seventy-three percent (188 cases) refer to transit systems under cost-plus reimbursement rules.

For the panel construction we resorted to two different informational sources. The starting database gathers information extracted from Federtrasporti annual reports concerning the years indicated above. From these reports we were able to derive the

46 As it occurs, on the contrary, in the Battese and Coelli (1992) model.

47 Federtrasporti (Rome) is a nationwide trade organization member of CISPEL (Confederazione delle Imprese di Servizi Pubblici degli Enti Locali) which associates the public-owned LPT companies and consortia in Italy. In 2001 it merged with Fenit (Federazione Nazionale Imprese di Trasporti), a nationwide trade organization which includes railway systems other than FS and private-owned bus operators, and assumed the new name ASSTRA. In 1998, the companies members of Federtrasporti was about 155, equal to 90% of the urban operators and to 50% of the extra-urban operators in Italy. The sample we utilize in our study may then be considered sufficiently representative.

48 Work is in progress to gather cost and technical information over the years 1993-1995 and 1999 also for the remaining 14 companies, so to have at disposal a seven-year balanced panel data, for a total of 315 observations.

49 Source: Ministero dei Trasporti e della Navigazione (1997).

main economic and productive data for each company in the sample, such as global production cost, labor cost, traveled kilometers, rolling stock size, average number of workers, and fuel consumption. Data were properly integrated by further information on cost, technical-environmental factors and type of regulatory scheme obtained through questionnaires sent to the companies with the support of Ceris-CNR and HERMES[50]. This additional investigation permitted to split the cost by productive factors other than labor, such as capital, fuel, and materials and services. Moreover, we retrieved relevant technical information (average load capacity and commercial speed of LPT vehicles, network size, average fleet age), in order to complement the data extracted from the Federtrasporti annual reports. To perform an analysis of the effects of regulatory schemes on x-efficiency, we need information which encompasses both the performance and the subsidization of the Italian public transit systems. To this end, in the questionnaire we included also a question on the reimbursement mechanism adopted by the competent local authority (Region, Province or Town Council)[51]. This rich source is probably unique in Italy as a tool of comparing regulatory systems to each other and over time.

5. Empirical results

ML estimates of the parameters of the model defined in [3]-[5] are given in Tables 1a and 1b. In particular, Table 1a reports the estimated coefficients, β, for the stochastic frontier cost function [3], while Table 1b presents the estimates of the inefficiency-related coefficients, δ, for the model [4] and the two variance parameters, γ and σ^2. The results are given in the second column of Tables 1a and 1b, indicated by *Full Model*. The third column of Tables 1a and 1b, indicated by *Restricted Model*, reports the ML estimates for the parameters of the "preferred" frontier model, to be discussed below, in which some coefficients in the general specification [3]-[5] are restricted to be zero.

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[50] Ceris-CNR is a division of the Italian National Research Council (CNR) which carries out economic research on firms and growth issues. HERMES (Turin) is a research center, created in July 2000, aimed at promoting economic research on transportation systems and other local public utilities.

[51] In particular, we asked the company to specify for each observed year if the subsidization was cost-plus (i.e., management contract) or fixed-price (i.e., net/gross cost contract) oriented. The answers were then checked by a direct telephone talk with the operators and a discussion with the juridical consultant of ASSTRA.

Table 1a. Maximum-likelihood estimates for parameters of the stochastic frontier cost function [3]*

Parameter	Full Model	Restricted Model
β_0	18.1137 (0.0315)	18.1246 (0.0293)
β_y	0.4741 (0.0465)	0.5123 (0.0152)
β_k	0.1319 (0.0977)	0.0469 (0.0302)
β_L	0.7873 (0.1295)	0.6917 (0.1144)
β_{MS}	0.1460 (0.0432)	0.1549 (0.0203)
β_{SP}	-0.1805 (0.0484)	-0.2352 (0.0372)
β_{Ly}	0.2785 (0.1389)	0.2321 (0.1159)
β_{Msy}	-0.0076 (0.0372)	0
β_{Lk}	-0.6100 (0.2986)	-0.5755 (0.2503)
β_{MSk}	-0.0300 (0.0750)	0
β_{yy}	-0.0524 (0.0696)	0
β_{kk}	-0.3892 (0.2330)	-0.2296 (0.0620)
β_{yk}	0.1450 (0.1063)	0.0503 (0.0146)
β_{LMS}	-0.0622 (0.2024)	0
β_{LL}	1.2522 (0.5990)	1.3902 (0.4730)
β_{MSMS}	0.2664 (0.0878)	0.1550 (0.0772)
β_{ySP}	0.1014 (0.5180)	0
β_{kSP}	-0.2683 (0.1016)	-0.0852 (0.0213)
β_{LSP}	-0.4193 (0.1602)	-0.2910 (0.1435)
β_{MSP}	0.3038 (0.0564)	0.2256 (0.0549)
β_{SPSP}	0.0836 (0.1226)	0
β_τ	-0.0157 (0.0048)	-0.0137 (0.0046)

* Estimated asymptotic standard errors are given in parentheses. All the independent variables except for time have been normalized to their sample mean value before the transformation in logarithms.

The signs of the first-order β parameters are all as expected, with the exception of the positive estimate for the quasi-fixed input coefficient, β_k. In fact, the evidence that the variable costs increase with larger rolling stocks is not consistent with the microeconomic theory[52]. With regards to this problem, an intense debate arose in the literature. According to Filippini (1996), the positive sign of K is due to a problem of multicolinearity in cases where there exists a positive correlation between dependent variable and capital indicator. The alternative argument suggested by Caves, Christensen, Tretheway and Windle (1985) and Windle (1988) is that the positive sign of K reflects an industry that does not minimize cost in the long term and therefore employs too much capital in the production process. This interpretation has been later proposed also in a study on the Italian urban transit systems carried out by Levaggi (1994). In this work the author argues that the inefficient use of capital could derive from the generous government programs of subsiding investments. This way of providing funds to purchase capital distorted the input allocation.

As far as the second-order β coefficients are concerned, it can be observed that the effects of the interaction of P_{MS} with Y, K and P_L, the interaction between Y and SP, and the quadratic terms for Y and SP, are very weak. Indeed, the estimated parameters, β_{MSy}, β_{MSk}, β_{LMS}, β_{ySP}, β_{yy} and β_{SPSP}, are all small and less than their estimated standard errors, hence asymptotic t-tests would lead to accept the null hypothesis of zero value for each of them[53].

The δ coefficients associated with the explanatory variables in the inefficiency model [5] are of particular interest to this study. It can be seen from the second column of Table 1b that the parameters have the right sign and are almost all statistically significant at the 1, 5 or 10 percent level of significance. The only exception is the estimate for the coefficient associated with the interaction between regulation dummy and time, $\delta_{R\tau}$. This parameter assumes a negative sign, which would indicate that the annual increase in x-inefficiency (highlighted by the positive sign of δ_τ) is less marked when firms face fixed-price regulatory schemes. However, its magnitude, -0.0308, is small compared to the estimated standard error, 0.0442, so the coefficient appears statistically insignificant. It is worthwhile to observe that the above interaction may also be interpreted as allowing the differential impact of regulation on cost inefficiency to depend on time. In this case, accepting the null hypothesis of zero value for $\delta_{R\tau}$ and rejecting that it is significantly less than zero would imply that the effects of regulatory

[52] This seems to be a general problem that characterizes the use of a variable cost model, not only in the transportation industry. For a discussion on these issues see also Fabbri (1998).

[53] Unless otherwise stated, all tests of hypothesis in this study are conducted at the 5% level of significance.

schemes are not statistically different across years, in particular the power of fixed-price contracts to reduce inefficiency does not become stronger with the passage of time.

Table 1b. Maximum-likelihood estimates for parameters of the stochastic cost inefficiency model [5]*

Parameter	Full Model	Restricted Model
δ_0	0.5881 (0.3532)	0.3546 (0.2098)
δ_R	-0.5544 (0.3751)	-0.6145 (0.2832)
δ_{SP}	-0.6475 (0.2786)	-0.6567 (0.1822)
δ_τ	0.0511 (0.0132)	0.0303 (0.0070)
δ_{RSP}	-1.0557 (0.7460)	-0.8592 (0.5192)
$\delta_{R\tau}$	-0.0308 (0.0442)	0
σ^2	0.0741 (0.0299)	0.0543 (0.0167)
γ	0.9561 (0.0198)	0.9334 (0.0263)
Log-likelihood	211.5864	205.6437

* Estimated asymptotic standard errors are given in parentheses. Variable *SP* has been normalized to its sample mean value and transformed in logarithms.

The remaining two parameters in Table 1b, $\sigma^2 \equiv (\sigma_v^2 + \sigma_u^2)$ and $\gamma \equiv \sigma_u^2/(\sigma_v^2 + \sigma_u^2)$, are associated with the variances of the random noise, v_{ft}, and the inefficiency term, u_{ft}. We note, in particular, that the ML estimate for γ is 0.9561 with asymptotic standard error of 0.0198, which indicates that the vast majority of residual variation is due to the x-inefficiency effects and these are likely to be highly significant in the analysis of the cost performance of the Italian LPT companies.

The hypotheses that the cost inefficiency effects are absent or that they have simpler distributions, together with the zero-restrictions on the β and δ parameters implied by the discussion above, have been statistically tested using the generalized likelihood-ratio (LR) test. The LR test statistics reported in Table 2 are calculated as

$$\Lambda = -2\{\ln[l(H_0)/l(H_1)]\} = -2\{\ln[l(H_0)] - \ln[l(H_1)]\}, \qquad [8]$$

where $l(H_0)$ and $l(H_1)$ are the values of the likelihood function under the null and alternative hypotheses, H_0 and H_1, respectively. If H_0 is true, this test statistic is usually assumed to be asymptotically distributed as a chi-square (χ^2) random variable with degrees of freedom equal to the number of restrictions involved. However, difficulties arise in testing hypotheses where γ is equal to 0 because $\gamma = 0$ lies on the boundary of the parameter space for γ, given that it cannot take negative values. In all these cases, if H_0 is true, the generalized LR statistic, Λ, has asymptotic distribution which is a mixture of chi-square distributions whose critical values are obtained from Table 1 in Kodde and Palm (1986)[54].

Table 2. Likelihood-ratio tests of hypotheses for parameters of the stochastic frontier cost function [3] and the cost inefficiency model [5]

Null hypothesis	Log-likelihood	χ^2-statistic	Decision
H_0: $\gamma = \delta_0 = \delta_R = \delta_{SP} = \delta_\tau = \delta_{RSP} = \delta_{R\tau} = 0$	179.258	64.655*	Reject H_0
H_0: $\gamma = \delta_0 = \delta_{SP} = \delta_\tau = 0$	182.847	57.478*	Reject H_0
H_0: $\delta_{R\tau} = 0$	211.513	0.147	Accept H_0
H_0: $\delta_{R\tau} = \beta_{MSy} = \beta_{MSk} = \beta_{yy} = \beta_{LMS} = \beta_{ySP} = \beta_{SPSP} = 0$	205.643	11.885	Accept H_0
Restriction: $\delta_{R\tau} = \beta_{MSy} = \beta_{MSk} = \beta_{yy} = \beta_{LMS} = \beta_{ySP} = \beta_{SPSP} = 0$			
H_0: $\gamma = \delta_0 = \delta_R = \delta_{SP} = \delta_\tau = \delta_{RSP} = 0$	171.718	67.849*	Reject H_0
H_0: $\gamma = \delta_0 = \delta_{SP} = \delta_\tau = 0$	176.310	58.665*	Reject H_0
H_0: $\delta_0 = \delta_R = \delta_{SP} = \delta_\tau = \delta_{RSP} = 0$	185.240	40.806	Reject H_0
H_0: $\delta_R = \delta_{SP} = \delta_\tau = \delta_{RSP} = 0$	192.671	25.943	Reject H_0
H_0: $\delta_0 = 0$	201.899	7.488	Reject H_0

* In this case the LR test statistic is asymptotically distributed as a mixture of chi-square distributions with degrees of freedom equal to the number of parameters assumed to be equal to zero in the null hypothesis H_0, provided H_0 is true. The critical values for this mixed χ^2-distribution are obtained from Table 1 in Kodde and Palm (1986).

It can be seen from Table 2 that the null hypothesis of absence of x-inefficiency effects from the model (i.e., H_0: $\gamma = \delta_0 = \delta_R = \delta_{SP} = \delta_\tau = \delta_{RSP} = \delta_{R\tau} = 0$) is strongly rejected at 1 per cent level of significance[55]. The second null hypothesis we consider, H_0: $\gamma = \delta_0 = \delta_{SP} = \delta_\tau = 0$, specifies that the inefficiency effects are not stochastic. If the parameter γ is zero, then the variance of the $u_{ft}s$ is zero and so the model reduces to a traditional mean response function in which the *z*-variables, R_{ft}, $(R_{ft} \times \ln SP_{ft})$ and $(R_{ft} \times \tau_{ft})$, are included in the cost function[56]. Once again, the H_0 hypothesis is rejected at

[54] For more on the use of this test in stochastic frontier models, see Coelli (1995) and Coelli and Battese (1996).

[55] The LR test statistic, 64.655, exceeds the 1% critical value for the mixed χ^2-distribution with 7 degrees of freedom, 17.755.

[56] Note that the parameters δ_0, δ_{SP} and δ_τ must be zero if γ is zero, given that the cost function already involves an intercept term, β_0, a first-order coefficient for the speed effect, β_{SP}, and a parameter associated with the year of observation, β_τ. If there are no random inefficiency effects in the model, then the coefficients δ_0, δ_{SP} and δ_τ are not identified.

the 1 percent level of significance[57]. Given the small estimate for the parameter $\delta_{R\tau}$ relative to its standard error underlined above, the third null hypothesis in Table 2 concerns the absence of significant effects on the cost inefficiency due to the interaction between time and regulation, $(R_{ft} \times \tau_{ft})$. As expected, H_0: $\delta_{R\tau} = 0$, i.e., the hypothesis that the marginal variation of the inefficiency term whit respect to time, $\partial u_{ft}/\partial \tau_{ft}$, does not depend on the reimbursement rule faced by the company (or, alternatively, the differential impact of fixed-price schemes, $\partial u_{ft}/\partial R_{ft}$, is substantially the same across years), is accepted[58]. Re-estimating the model without $\delta_{R\tau}$, the estimates of the other parameters, β and δ, were little different from those obtained for the more general model, but the coefficients associated with the interaction of P_{MS} with Y, K and P_L, the interaction between Y and SP, and the quadratic terms for Y and SP persisted to be small and less than their estimated standard errors. Indeed, the LR statistic reported in Table 2 for testing the joint hypothesis H_0: $\delta_{R\tau} = \beta_{MSy} = \beta_{MSk} = \beta_{yy} = \beta_{LMS} = \beta_{ySP} = \beta_{SPSP} = 0$ is not significant[59] and so we consider that the preferred stochastic frontier model has the seven parameters, $\delta_{R\tau}$, β_{MSy}, β_{MSk}, β_{yy}, β_{LMS}, β_{ySP} and β_{SPSP}, constrained to be equal to zero.

The ML estimates for the parameters of the restricted model are presented in the third column of Tables 1a and 1b. It can be seen that all the β and δ coefficients for this model are larger than their estimated standard errors and most of them are statistically significant at the 1 percent level[60]. Table 2 reports the LR statistics for testing the null hypotheses of absence of inefficiency effects (sixth row) and of absence of stochastic effects (seventh row). Both values are not significant[61]. Similarly, the null hypotheses that the $u_{ft}s$ are altogether unrelated to the z-variables (eighth row), that they are not a linear function of the subsidization mechanisms, the network commercial speed, the year of observation and the interaction between regulation and speed (ninth row), and that they do not include an intercept parameter (tenth row), are all also rejected at the

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57 In this case the LR test statistic, 57.478, exceeds the 1% critical value for the mixed χ^2-distribution with 4 degrees of freedom, 12.483.

58 The value of the χ^2-statistic reported in Table 2, 0.147, is less than the 1%, 5% and 10% critical values for the χ^2-distribution with 1 degree of freedom, which are 6.634, 3.841 and 2.705, respectively.

59 The value of the χ^2-statistic, 11.885, is less than the upper 10 % point for the χ^2-distribution with 7 degrees of freedom, 12.017. Thus the null hypothesis is accepted at the 1, 5 and 10 percent levels of significance.

60 The null hypothesis of zero value is rejected at the 1% level of significance (by asymptotic t-tests) only for the coefficient associated with the quasi-fixed input, β_k, which is statistically significant at the 10% level, and for the parameters β_{Ly}, β_{Lk}, β_{LSP} and β_{MSMS} in the frontier cost function, and the parameters δ_0, δ_R and δ_{RSP} in the cost inefficiency model, which are all statistically significant at the 5% level.

61 In the first case, the LR test statistic, 67.849, exceeds the 1% critical value for the mixed χ^2-distribution with 6 degrees of freedom, 16.074, while in the second case, the LR test statistic, 58.665, exceeds the 1% critical value for the mixed χ^2-distribution with 4 degrees of freedom, 12.483.

1% level of significance[62]. On the whole, these results show that the present data support the presence of stochastic x-inefficiency effects. Furthermore, the model specified for the inefficiency effects, involving a constant term, regulatory schemes, network speed, year of observation and interaction between speed and regulation type, appears to be a significant component in our stochastic cost frontier analysis of the Italian LPT industry.

In the following Section (5.1) we will briefly take a look at the estimates of frontier cost elasticities and the technological properties for the preferred translog model. We postpone to Section 5.2 the discussion concerning detected x-inefficiencies and the effects of regulatory schemes, which are our primary interest in this study.

5.1. Technical characteristics

Since all the variables (excepting time) in the cost function have been normalized to their sample mean value, and the variable cost as well as the regressors are in natural logarithm, the estimated first-order coefficients in the third column of Table 1a can be usefully interpreted as frontier cost elasticities for the *average operator* of the industry[63]. The focus of the analysis, in particular, is on the elasticities with respect to output, β_y, capital stock, β_k, commercial speed, β_{SP}, and time, β_τ. These have been utilized to infer the characteristics of technology (evaluated at the sample mean) presented in Table 3, which highlights the separated effects on frontier costs attributable to short-run (*SRS*) and long-run returns to scale (*LRS*), commercial speed improvements, and Hicks neutral technological change.

The analysis, revealing the presence of short- and long-run scale economies, validates the empirical evidence emerged in our previous study on the Italian LPT technology. Indeed, asymptotic t-tests lead to accept both hypotheses that *SRS* and *LRS* are significantly greater than one[64]. The estimated *SRS*, 1.95, show that, given the endowment of quasi-fixed input, a more than proportional (almost double) output growth could be achieved by a proportional increase in the use of all variable factors, allowing the operator to reduce its unitary cost of production. This feature is

62 The upper 1% points for the χ^2-distribution with 5, 4 and 1 degrees of freedom are 15.086, 13.276 and 6.634, respectively.

63 The *average operator* (the point of normalization) corresponds to an hypothetical firm operating at an average level of production (542,216 millions of places*traveled-kms), using an average stock of quasi-fixed input (276 age-vehicles), and facing average variable input prices (P_L = 68.3 millions lire/workers; P_{MS} = 13.41 lire/seat-km; P_F = 1,037 lire/litre) and average commercial speed (23.3 kms/h) over the sample.

64 Test statistics for the estimated *SRS* and *LRS* are 16.440 and 29.771, respectively. Both values are remarkably larger than the one-sided critical value of 2.326 for the standard normal distribution at the 1 percent level of significance.

emphasized on Figure 1, which represents the estimated average variable costs (AVC) and marginal costs (MC) for the sample mean company in 1996[65] and highlights a trend decreasing with the level of output[66]. So far as *LRS* are concerned[67], the estimate, 1.86, implies a sub-optimal scale with respect to the long-run equilibrium. On the whole, these results highlights the existence of unused capacity and support the conjecture that local monopoly is the relevant organization in the industry, at least for medium-sized firms[68].

Table 3. Technology characteristics evaluated at the mean of the data (average firm)*

Returns to scale Short-run $[1/\beta_y]$	Returns to scale Long-run $[(1-\beta_k)/\beta_y]$	Elasticity with respect To commercial speed $[\beta_{SP}]$	Rate of Hicks neutral technological change $[-\beta_\tau]$
1.9519 (0.0579)	1.8604 (0.0289)	-0.2352 (0.0372)	0.0137 (0.0046)

* Estimated asymptotic standard errors are given in parentheses.

The estimated frontier cost elasticity with respect to the average speed of the network in Table 3 bears out our insights about the influence on the production process of the specific environmental conditions characterizing the area where the service is provided. Increasing speed of LPT vehicles by 10 percent permits to reduce the level of operating costs for the average firm by about 2.3 percent. This result underlines the importance of appropriate public policies concerning the local traffic regulation.

65 Year 1996 represents the average period of observation in our sample.

66 To predict the trend of AVC and MC presented in Figure 1, only the level of output has been let to vary across the sample (from the smallest value, 33 millions of places*travelled-kms, up to 1,409,919 millions), all the other variables (input prices, capital stock and commercial speed) remaining unchanged at the average firm values.

67 We can evaluate the long-run returns to scale by applying the algorithm first suggested by Caves et al. (1981) and indicated in Table 3.

68 In fact, short- and long-run scale economies have been calculated at all production levels of the sample (with the other variables fixed at the average firm values) and increasing *SRS* and *LRS* are observed at all data points. Since the cost elasticity with respect to output does not depend on the starting production level in the preferred frontier model (the quadratic term for output, β_{yy}, is specified to be zero), the value of *SRS* is constant throughout the sample and equal to 1.95. The estimate of *LRS*, on the contrary, depends on the output through the cost elasticity with respect to capital (in which β_{yk} = 0.05) and decreases from 2.81 for the lowest production level (corresponding to Sondrio ASM in 1993) to 1.57 for the highest production level in the sample (corresponding to Torino ATM in 1994).

Figure 1. Estimated average variable costs (AVC) and marginal costs (MC)

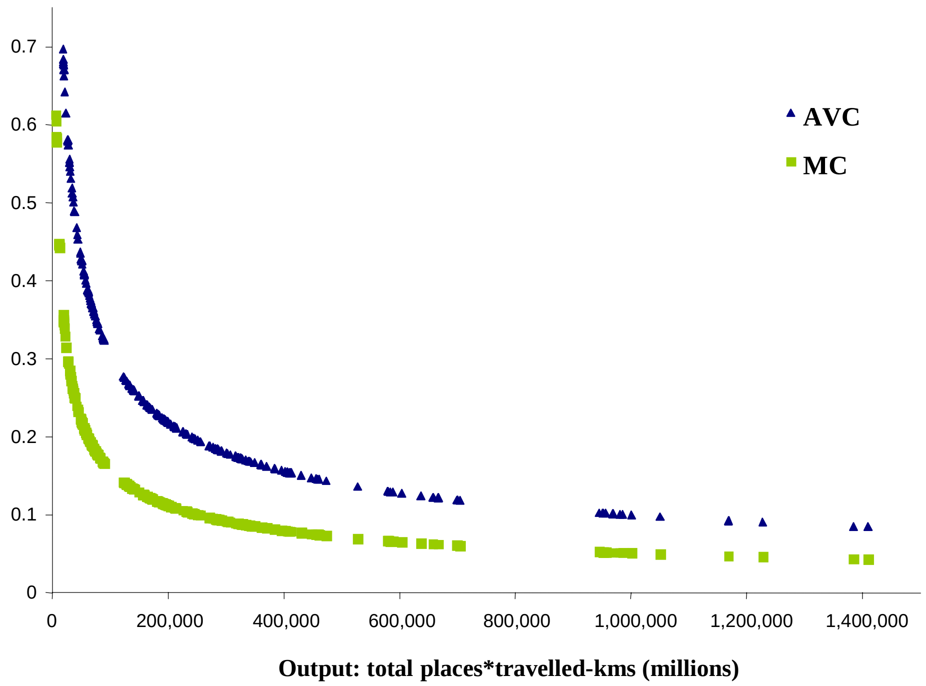

Table 3 also presents the estimated rate of Hicks neutral technological change, i.e., the rate of cost diminution over years 1993 to 1999 ($-\partial \ln VC_{ft} / \partial \tau_{ft} = -\beta_{\tau}$)[69]. As expected, variable costs are negatively related to the time variable: other things remaining unchanged, the annual rate of cost reduction due to technical progress is about 1.4 percent. This decrease in costs over time presumably reflects, to some extent, the greater care of the road maintenance and the replacement of worn-out fleet and the introduction of more fuel-efficient models of vehicles[70], made possible by the generous grants-in-aid government programs.

5.2. Cost inefficiency and effects of regulatory schemes

From Table 1b, we note first that the ML estimate for γ in the preferred model is 0.9334 with asymptotic standard error of 0.0263. This result is consistent with the conclusion that the true γ-value is accepted to be greater than zero (in the LR tests above) and the traditional average response function is not an adequate representation of the data. However, although the vast majority of residual variation is due to the cost

So, the above insights can also be extended to both the small- and large-sized companies.

69 See Caves et al. (1981).

70 Between 1993 and 1996 the fuel efficiency (kilometer run per liter of fuel) of the LPT companies included in our sample has increased on average from 2.5 to 2.8. Furthermore, the total expenditure for spares and repairs has decreased on average from 9,919,257 lire per vehicle in 1996 to 9,765,244 lire per vehicle in 1999. This probably contributed to the cost reduction over time highlighted above.

inefficiency effects, u_{ft}, we also see that the γ estimate is significantly less than one[71], to indicate that our stochastic frontier model [3]-[5] may be significantly different from a deterministic frontier specification, in which there are no random errors, v_{ft}, in the cost function.

The predicted x-inefficiencies for each one of the 45 transit companies over the different years involved are presented in Table 4. These estimates refer to the expression defined by equation [7] and have been obtained using the predictor presented in equation [A.20] of the Appendix. The mean overall cost inefficiency, corresponding to square "all firms-all years", is found to be 1.137[72]. This means that, on average, the cost of production exceeds the minimum level frontier by 13.7 percent because of x-inefficiency. The positive coefficient for τ_{ft} in Table 1b (δ_τ = 0.0303) suggests that the inefficiencies of the Italian LPT firms tended to increase throughout the seven-year period. First row of Table 4, which reports the estimates for mean cost inefficiency over time, confirms this tendency to worsen the performance: on average, the level of x-inefficiency has increased slightly, from 12.2 percent in 1993 to 14.2 in 1999, with an upward swing during 1993-1995 and 1998-1999 and a brief downward swing over the period 1996 to 1997[73]. As mentioned in Section 3.2, one can possibly trace the deterioration of cost efficiency during the first half of the nineties in the laxity induced by the several actions taken by the Government with the purpose of covering the old deficits of LPT companies through extraordinary funds. On the contrary, the temporary efficiency recovery during 1996-1997 could be linked to expectations of more tight financial constraints triggered by the promulgation of the reform Law n. 549 in 1995, whereas the new rise in x-inefficiency observed in the years 1998 and 1999 probably reflects a let-up in the managerial effort induced by the delay in implementing the reform.

Although there is a general increase in the x-inefficiency of the transit companies over time, Table 4 shows that the individual predicted values vary considerably among firms in each year and they also change up and down over time for a given company. This leads to investigate the role played by the other *z*-factors included in model [5] that, jointly with time, determine such a variability in the inefficiency levels.

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[71] The test statistic is -2.532, that is larger (in absolute value) than the one-sided critical value of -2.326 for the standard normal distribution at the 1 percent level of significance.

[72] This value is calculated as the arithmetic average of the predictors for the individual cost inefficiency of the sample firms over all the observations involved.

[73] We also computed the average annual rate of variation in the level of cost inefficiency. This is equal to about +3%, which is consistent with the estimate for the parameter associated to the year of observation in the inefficiency model [5], δ_τ.

Table 4. Estimated cost inefficiency by firm and year

LPT company	Year							
	1993	1994	1995	1996	1997	1998	1999	All Years
All Firms	1.122	1.144	1.148	1.137	1.131	1.135	1.142	1.137
Firm 1	-	-	-	1.315	1.343	1.376	-	1.344
Firm 2	1.066	1.132	1.141	1.080	1.135	1.126	1.150	1.118
Firm 3	-	-	-	1.024	1.025	1.029	-	1.026
Firm 4	1.036	1.032	1.034	1.035	1.057	1.065	1.039	1.042
Firm 5	1.189	1.180	1.097	1.036	1.034	1.031	1.028	1.085
Firm 6	1.096	1.059	1.097	1.052	1.042	1.045	1.064	1.065
Firm 7	1.230	1.237	1.203	1.168	1.210	1.250	1.244	1.220
Firm 8	-	-	-	1.048	1.059	1.081	-	1.062
Firm 9	-	-	-	1.203	1.197	1.129	-	1.176
Firm 10	1.069	1.221	1.339	1.349	1.272	1.311	1.224	1.255
Firm 11	1.245	1.276	1.179	1.115	1.122	1.173	1.212	1.188
Firm 12	1.039	1.039	1.075	1.082	1.085	1.083	1.117	1.074
Firm 13	-	-	-	1.305	1.299	1.287	-	1.297
Firm 14	1.042	1.045	1.042	1.051	1.054	1.062	1.057	1.050
Firm 15	-	-	-	1.052	1.041	1.046	-	1.046
Firm 16	1.063	1.178	1.223	1.401	1.259	1.470	1.741	1.333
Firm 17	1.074	1.063	1.047	1.045	1.046	1.020	1.015	1.044
Firm 18	-	-	-	1.054	1.060	1.050	-	1.054
Firm 19	-	-	-	1.203	1.284	1.249	-	1.245
Firm 20	1.117	1.130	1.098	1.068	1.066	1.044	1.036	1.080
Firm 21	1.121	1.146	1.055	1.076	1.092	1.103	1.097	1.098
Firm 22	-	-	-	1.108	1.095	1.112	-	1.105
Firm 23	1.024	1.053	1.036	1.078	1.063	1.040	1.024	1.045
Firm 24	1.170	1.152	1.245	1.141	1.143	1.113	1.211	1.167
Firm 25	1.724	1.771	1.802	1.834	1.717	1.715	1.586	1.735
Firm 26	-	-	-	1.017	1.043	1.050	-	1.036
Firm 27	1.049	1.053	1.085	1.042	1.050	1.046	1.081	1.059
Firm 28	1.050	1.038	1.039	1.061	1.046	1.046	1.072	1.050
Firm 29	-	-	-	1.052	1.053	1.059	-	1.054
Firm 30	1.132	1.142	1.146	1.174	1.106	1.140	1.116	1.136
Firm 31	1.122	1.169	1.131	1.123	1.090	1.076	1.057	1.109
Firm 32	-	-	-	1.078	1.104	1.188	-	1.123
Firm 33	1.065	1.089	1.063	1.026	1.027	1.025	1.035	1.047
Firm 34	1.045	1.065	1.056	1.130	1.116	1.115	1.133	1.094
Firm 35	1.045	1.041	1.093	1.200	1.109	1.111	1.073	1.096
Firm 36	1.045	1.064	1.038	1.059	1.050	1.062	1.054	1.053
Firm 37	1.192	1.299	1.415	1.134	1.324	1.071	1.102	1.219
Firm 38	1.254	1.284	1.295	1.216	1.215	1.249	1.250	1.251
Firm 39	1.137	1.149	1.149	1.225	1.220	1.213	1.250	1.191
Firm 40	-	-	-	1.196	1.023	1.075	-	1.098
Firm 41	1.041	1.061	1.042	1.064	1.047	1.045	1.046	1.049
Firm 42	1.251	1.230	1.266	1.401	1.383	1.345	1.267	1.306
Firm 43	1.044	1.036	1.043	1.025	1.023	1.025	1.024	1.031
Firm 44	1.028	1.030	1.031	1.020	1.021	1.021	1.020	1.024
Firm 45	-	-	-	1.054	1.070	1.087	-	1.070

Our primary concern, in this work, is with the differential impact of regulatory schemes on cost efficiency. From Table 1b, the negative sign of δ_R (-0.6145), the parameter related to the subsidization mechanisms as such, without their interaction with network characteristics, seems to back our opening conjecture of lower x-inefficiency levels for the units run under fixed-price schemes. Indeed, when compared over time, the results of Table 4 indicate a tendency of predicted cost inefficiency to diminish for most of companies facing a transition from cost-plus to fixed-price reimbursement mechanisms. The differential impact of regulation is clearly observable in many cases where the subsidization practice changed from 1996 onwards, such as, for instance, Firm 5 (9.7 percent in 1995, 3.6 in 1996 and 3.4 percent in 1997) or Firm 43 (4.3 percent in 1995, 2.5 in 1996 and 2.3 percent in 1997). A similar evidence is found for the transit systems which shifted from a cost-plus to a fixed-price scheme the following year or three years later, as it occurred for Firm 41 (6.4 percent in 1996, 4.7 in 1997 and 4.5 percent in 1998), Firm 4 (6.5 percent in 1998, 3.9 percent in 1999), or Firm 30 (14.0 percent in 1998, 11.6 percent in 1999)[74].

It is worthwhile to highlight that the magnitude of the efficiency recovery differs from case to case, and not all the firms which faced a regulatory change exhibit better performances after the transition. This is due to the fact that the inefficiency estimates reported in Table 4 represent the combined effect of the regulation dummy and two other explanatory variables (besides time), viz., the average commercial speed of vehicles, that is a proxy for network characteristics, and its interaction with the subsidization mechanisms. Table 1b shows that an increase in the network speed tends to lower x-inefficiency (δ_{SP} = -0.6567), as the transit company faces more favourable exogenous operating conditions, and this effect is strongest for the units subjected to fixed-price schemes (δ_{RSP} = -0.8592), presumably because of the higher cost reducing effort exerted by managers under this type of regulation. As explained in Section 3.2, from the latter result it is also proper to infer that when the intrinsic inefficiency of a network is too high (here due to a very low commercial speed), the impact of regulatory constraints on the overall cost efficiency becomes modest and in the extreme circumstances is no longer perceptible (Gagnepain and Ivaldi, 1998). Thus the greater efficiency recovery for some of the companies moved towards fixed-priced mechanisms can be partially attributed to better network characteristics as reflected in the higher

[74] Other situations of companies in which the introduction of fixed-price schemes generated a significant fall in the level of x-inefficiency are represented by Firm 27 (8.5 percent in 1995, 4.2 in 1996 and 5.0 percent in 1997), Firm 33 (6.3 percent in 1995, 2.6 in 1996 and 2.7 percent in 1997) and Firm 44 (3.1 percent in 1995, 2.0 in 1996 and 2.1 percent in 1997), so far as the transition in 1996 is concerned, and Firm 36 (6.2 percent in 1998, 5.4 percent in 1999) and Firm 20 (4.4 percent in 1998, 3.6 percent in 1999), as regards the transition in 1999.

level of average commercial speed[75]. On the other hand, the modest effects of the regulatory change detected for some units in the sample, or even deterioration in the performance showed by others, is possibly imputable to worsened operating conditions[76], i.e., lower network speed, in addition to the impact exerted by time and the exogenous shocks captured by the stochastic term w_{ft} that can have unfavorable repercussions on cost efficiency.

So far we have focused on the differential impact of regulatory schemes over time, by comparing predicted inefficiencies for a given LPT company before and after the introduction of a fixed-price mechanism. To better highlight the separated effects on cost efficiency of the regulation and network characteristics, as well as the relevance of their mutual interaction, it may be convenient to fix the attention on cross-sectional comparisons among firms. In fact, it is in this context that the most marked variability in both the subsidization rules and the levels of average commercial speed is observable. In such a way, we are also able to exploit the inefficiency estimates concerning companies for which a time-serial match of the two types of regulation is not possible, because of the lack of information over the period leading up the adoption of fixed-price schemes (e.g., Firm 3, Firm 8, Firm 18).

To our end, we concentrate on the individual predicted inefficiencies pertaining to years 1996, 1997 and 1998. We have chosen this sub-period since, excepting Firm 41, for which a fixed-price oriented scheme is in force from 1997 only, the other 43 companies of our sample are univocally characterized by a definite regulatory mechanism during this years. In order to leave aside time effects, we calculated a mean inefficiency level over the period for each operator and considered the x-inefficiency values as average realizations of a specific subsidization rule. This allows us to classify the LPT firms on the basis of their inefficiency levels and to interpret the ensuing ranking in terms of the different regulatory schemes and network characteristics faced by each company. The list is presented in Table 5. Instead of reporting $C\hat{I}_{ft}$, we computed the percentage increase in costs due to x-inefficiency from the expression $\{C\hat{I}_{ft}-1\}$, so the entries in Table 5 can be directly taken as mean cost distortions over the frontier between 1996 and 1998. They have been ranked from the best performance (Firm 44), characterized by observed operating costs that are, on average, 2.07 percent only above the frontier, to the worst performance (Firm 25), for which the cost

[75] It is the case, for instance, of Firm 43, which reduced its level of x-inefficiency by about 42% between 1995 and 1996, or Firm 33, for which the recovery during the same period reached about 60%.

[76] Firm 14 and Firm 35, for example, faced a decline in their network speed level between 1995 and 1996; while during the same period their cost inefficiency increased by about 21% and 115%, respectively, despite the transition to fixed-price reimbursement schemes. The performance deterioration exhibited by Firm 35 is drastic because of the already poor level of commercial speed faced in 1995 (16.5 kms/h).

distortion climbs to more than 70 percent. An important result emerging from Table 5 is that 8 of the top 11 firms were subjected to fixed-price mechanisms[77], whereas 9 of the bottom 11 companies faced a cost-plus regulation[78]. Once again, our findings tend to corroborate the theoretical argument about the efficacy of high powered incentive schemes in increasing efficiency.

Table 5. Ranking of firms by mean cost distortion over the frontier (time period 1996-1998)*

LPT company	Cost distortion		LPT company	Cost distortion	
Firm 44	0.0207	(0.000)	Firm 31	0.0964	(0.024)
Firm 43	0.0243	(0.001)	Firm 40	0.0978	(0.089)
Firm 3	0.0258	(0.003)	Firm 22	0.1052	(0.009)
Firm 33	0.0261	(0.001)	Firm 2	0.1137	(0.030)
Firm 5	0.0336	(0.002)	Firm 34	0.1202	(0.008)
Firm 26	0.0367	(0.017)	Firm 32	0.1232	(0.057)
Firm 17	0.0368	(0.015)	Firm 24	0.1324	(0.017)
Firm 15	0.0462	(0.006)	Firm 11	0.1366	(0.032)
Firm 6	0.0463	(0.005)	Firm 35	0.1401	(0.052)
Firm 27	0.0492	(0.010)	Firm 30	0.1401	(0.034)
Firm 28	0.0513	(0.008)	Firm 9	0.1761	(0.041)
Firm 41	0.0519	(0.010)	Firm 37	0.1764	(0.132)
Firm 4	0.0522	(0.015)	Firm 7	0.2096	(0.041)
Firm 29	0.0545	(0.004)	Firm 39	0.2195	(0.006)
Firm 18	0.0546	(0.005)	Firm 38	0.2269	(0.019)
Firm 14	0.0556	(0.005)	Firm 19	0.2455	(0.041)
Firm 36	0.0576	(0.006)	Firm 13	0.2966	(0.009)
Firm 20	0.0594	(0.014)	Firm 10	0.3105	(0.035)
Firm 23	0.0606	(0.019)	Firm 1	0.3450	(0.031)
Firm 8	0.0625	(0.017)	Firm 42	0.3763	(0.028)
Firm 45	0.0703	(0.016)	Firm 16	0.3769	(0.108)
Firm 12	0.0836	(0.002)	Firm 25	0.7555	(0.068)
Firm 21	0.0906	(0.014)			

* Standard deviations are given in parentheses.

To better understand the role played by network characteristics (i.e., average commercial speed) in the above ranking, we assigned the individual predicted inefficiencies to four speed classes, defined in terms of brackets of average kilometers to the hour: very low speed, $SP_{vl} \in [13, 17.3]$; low speed, $SP_l \in [17.4, 23.2]$; high speed, $SP_h \in [23.3, 31.4]$; very high speed, $SP_{vh} \in [31.5, 45.5]$. We distinguished then operators

[77] These are: Firm 44, Firm 43, Firm 3, Firm 33, Firm 5, Firm 6, Firm 27, and Firm 28.
[78] These are: Firm 25, Firm 16, Firm 1, Firm 13, Firm 19, Firm 38, Firm 39, Firm 7, and Firm 37.

subjected to cost-plus regimes, 0, from units run under fixed-price ones, 1. Finally, we crossed the two types of categories, speed class and subsidization mechanism, and for each of resultant groups we computed a mean cost distortion over the frontier. These values are reported in Table 6 (shadowed square), that also presents mean cost distortions by regulatory scheme and speed class regardless of their interaction (first row and first column, respectively), together with the percentage decrease in x-efficiency attainable by shifting from cost-plus to fixed-price regimes (*regulation effect*) and/or by improving operating conditions of the network (*speed effect*).

Table 6. Mean cost distortion over the frontier by regulatory scheme and average commercial speed class (time period 1996-1998)

Average commercial speed class	Subsidization mechanism			
	All Schemes	Cost-plus scheme [0]	Fixed-price scheme [1]	*Regulation effect : (1 – 0)/0*
All speed classes*	0.1349	0.1604	0.0916	-42.89%
Very low speed [SP_{vl}]	0.1791	0.2076	0.1593	-23.26%
Low speed [SP_l]	0.1651	0.2058	0.0901	-56.22%
High speed [SP_h]	0.0983	0.1141	0.0432	-62.14%
Very high speed [SP_{vh}]	0.0692	0.1092	0.0242	-77.84%
Speed effect :				
$(SP_l – SP_{vl})/SP_{vl}$	-7.82%	-0.87%	-39.65%	
$(SP_h – SP_l)/SP_l$	-40.46%	-44.56%	-52.05%	
$(SP_{vh} – SP_h)/SP_h$	-29.60%	-4.29%	-43.98%	

* Commercial speed classes have been defined in terms of brackets of average kilometers to the hour: $SP_{vl} \in [13, 17.3]$; $SP_l \in [17.4, 23.2]$; $SP_h \in [23.3, 31.4]$; $SP_{vh} \in [31.5, 45.5]$.

First of all, the entries in Table 6 clearly confirm that both network characteristics and regulatory constraints matter in determining x-efficiency of LPT firms: for a company facing medium levels of commercial speed, the introduction of high powered incentive schemes allows, on average, an efficiency recovery around 43% (first row-last column); similarly, more favorable traffic conditions for the LPT vehicles imply lower cost inefficiencies, with reductions which range from about 8% up to 40% according to the starting level of network speed, SP_{vl}, SP_l, SP_h, or SP_{vh} (first column-bottom three rows).

Second, it emerges a general tendency of the regulation effect to become stronger as we move towards higher speed classes; on the other hand, the average efficiency gain that can be realized by making mobility more flowing are always greater in case of fixed-price regimes, given the higher effort level exerted by managers in the allocation of productive resources. In particular, two opposite groups of operators (the ones more shadowed in the square of Table 6) are worth to be highlighted: in the first one, which includes units characterized by very high levels of network speed, the favorable operating conditions combined with fixed-price regulations leads to remarkable inefficiency decrease, on average around 78% (from 10.92 to 2.42 percent); in the second group, gathering very slow speed networks, since the exogenous technical efficiency is likely to be rather low, the more intensive effort activity provided by managers in case of fixed-price schemes has a moderate effect on the x-inefficiency (-23%) and the global cost distortions over the frontier remain heavy, on average about 16 percent.

These results can help explain the ranking of companies presented in Table 5 above. It should be more clear why the top four positions are held by companies facing very high levels of commercial speed combined with incentive subsidization mechanisms, while at the bottom of the list one observes mainly firms under cost-plus regulation with a very slow network speed. At the same time, we are also able to account for both the presence of companies subjected to cost-plus subsidization among the top ten performances, due to the favorable characteristics of their network[79], and the positioning of operators constrained by fixed-price schemes among the worst ten positions, because of the very low levels of their commercial speed[80].

6. Conclusion and policy implications

On the whole, the results of this exploratory study indicate a significant impact of regulatory constraints on the cost efficiency of the Italian LPT companies. First, the fact that fixed-price schemes provide more incentives for efficiency is validated: given similar network characteristics, operators run under a fixed-price mechanism have a lower cost distortion than operators subjected to a cost-plus regulation. Moreover, to

[79] It is the case, for instance, of operators such as Firm 26 or Firm 17, which face average commercial speeds equal to around 45 and 27 kms/h, respectively.

[80] In this situation, we find Firm 10 (16 kms/h) and Firm 48 (15 kms/h).

some extent the inefficiency differentials among companies can be due to differences in the commercial speed levels. These latter contribute to determine what Gagnepain & Ivaldi (1998) called intrinsic inefficiency of a network and can seriously undermine the efficacy of incentive regulatory policies. In view of the evidence found for the operators facing very low commercial speed levels, if the exogenous operating conditions become too unfavorable, then fixed-price subsidization mechanisms are less successful instruments for recovering efficiency.

These findings provide useful guidelines for the policy interventions concerning local mobility. Significant reductions of x-inefficiency can be achieved by introducing fixed-price schemes of subsidization, and the ongoing reform correctly moves towards this direction. A proper definition of quality and cost standards is requested, so that the service contract between local authority and LPT operator gives the manager the incentives to optimize the allocation of productive resources. Our results also underline the impact of network characteristics and confirm the importance of local traffic regulation, already stressed by Fraquelli et al. (2001b). In fact, a more flowing mobility for LPT vehicles would have positive effects on both the technology (higher commercial speeds lower the minimum-cost frontier) and the x-efficiency levels (higher commercial speeds move firm performances closer to the best-practice behavior). This could be pursued, for instance, by acting on factors such as the re-allocation of existing road space away from private vehicles towards public passenger transport (e.g., reserved lanes for trams and buses, restrictions on parking and traffic of cars and taxis), or the provision of incentives for the use of public modes (e.g., good intra- and inter-modal timetable coordination, introduction of multi-modal travelcards).

In conclusion, there is a scope for transport policy to increase cost efficiency of Italian LPT companies. Efforts have to be intensified in the twofold direction of replacing cost-plus subsidization mechanisms with high-powered incentive schemes as well as improving exogenous operating conditions of the network. Indeed, a peculiarity of our study is to highlight the complementarity between the effects exerted by these two instruments. Local authorities will have to define the proper mix of interventions according to the specific regulatory framework and environmental factors faced by single LPT firms.

Appendix

The log-likelihood function presented in the Appendix of the Battese and Coelli (1993) working paper refers to a stochastic frontier *production* function, with the $u_{ft}s$ interpreted as pure *technical inefficiency* effects, which cause the firm to operate below the production frontier. If we wish to specify a stochastic frontier *cost* function, we have to alter the global error term specification from $\psi_{ft} = (v_{ft} - u_{ft})$, as in Battese and Coelli (1993), to $\psi_{ft} = (v_{ft} + u_{ft})$, as in the equation [2] reported in the text[81]. The $u_{ft}s$ now define how far the firm operates above the cost frontier and involve both *technical and allocative inefficiencies*. The log-likelihood function for the cost frontier specification analogue of the Battese and Coelli model can be obtained by making a few simple sign changes and is reproduced here.

For simplicity of presentation of results in this Appendix, we assume that the stochastic frontier cost model [1]-[2] is expressed by

$$vc_{ft} = vc(x_{ft}; \beta) + \psi_{ft} \qquad [A.1]$$

$$\psi_{ft} = v_{ft} + u_{ft} \qquad [A.2]$$

$f = 1, \ldots, F$, and $t = 1, \ldots, T_f$,

where $vc_{ft} = \ln VC_{ft}$, $vc(.) = \ln VC(.)$ and x_{ft} is a vector which groups the arguments of the variable cost function, Y_{ft}, P_{ft}, Z_{ft}, and τ_{ft}; further, $v_{ft} \sim i.i.d.\ N(0, \sigma_v^2)$ and $u_{ft} \sim N^+(\delta' z_{ft}, \sigma_u^2)$.

The density functions for v_{ft} and u_{ft} are

$$f_V(v) = \frac{\exp\left\{-\frac{1}{2} v^2 / \sigma_v^2\right\}}{\sqrt{2\pi}\sigma_v}, \qquad -\infty < v < \infty \qquad [A.3]$$

and

$$f_U(u) = \frac{\exp\left\{-\frac{1}{2} (u - \delta' z)^2 / \sigma_u^2\right\}}{\sqrt{2\pi}\sigma_u \Phi[\delta' z / \sigma_u]}, \quad u \geq 0, \qquad [A.4]$$

where the subscripts, f and t, are omitted for convenience in the presentation; and $\Phi[\cdot]$ is the standard normal cumulative distribution function.

[81] The inefficiency effect, u_{ft}, is *added* in the stochastic cost frontier instead of being *subtracted*, as in the case of the stochastic production frontier, because the cost function represents *minimum cost*, whereas the production function represents *maximum output*.

Given the independence assumption, the joint density function for u and v is the product of their individual density functions, and so

$$f_{U,V}(u,v) = \frac{\exp-\frac{1}{2}\{[(u-\delta' z)^2/\sigma_u^2]+v^2/\sigma_v^2\}}{2\pi\sigma_u\sigma_v\Phi[\delta' z/\sigma_u]}; \qquad \text{[A.5]}$$

since $\psi = v + u$, the joint density function for ψ and u is

$$f_{\Psi,U}(\psi,u) = \frac{\exp-\frac{1}{2}\{[(\psi-u)^2/\sigma_v^2]+[(u-\delta' z)^2/\sigma_u^2]\}}{2\pi\sigma_u\sigma_v\Phi[\delta' z/\sigma_u]}, \quad u \geq 0$$

$$= \frac{\exp-\frac{1}{2}\{[(u-\mu_*)^2/\sigma_*^2]+(\psi^2/\sigma_v^2)+(\delta' z/\sigma_u)^2-(\mu_*/\sigma_*)^2\}}{2\pi\sigma_u\sigma_v\Phi[\delta' z/\sigma_u]}, \quad \text{[A.6a]}$$

or, alternatively,

$$f_{\Psi,U}(\psi,u) = \frac{\exp-\frac{1}{2}\{[(u-\mu_*)^2/\sigma_*^2]+[(\psi-\delta' z)^2/(\sigma_v^2+\sigma_u^2)]\}}{2\pi\sigma_u\sigma_v\Phi[\delta' z/\sigma_u]}, \qquad \text{[A.6b]}$$

where

$$\mu_* = \frac{\sigma_v^2\delta' z+\sigma_u^2\psi}{\sigma_v^2+\sigma_u^2} \qquad \text{[A.7]}$$

and

$$\sigma_*^2 = \sigma_v^2\sigma_u^2/(\sigma_v^2+\sigma_u^2). \qquad \text{[A.8]}$$

Thus the marginal density function for $\psi = v + u$ is obtained by integrating u out of $f_{\Psi,U}(\psi,u)$, which yields

$$f_\Psi(\psi) = \frac{\exp-\frac{1}{2}\{(\psi^2/\sigma_v^2)+(\delta' z/\sigma_u)^2-(\mu_*/\sigma_*)^2\}}{\sqrt{2\pi}\sigma_u\sigma_v\Phi[\delta' z/\sigma_u]}\int_0^\infty \frac{\exp-\frac{1}{2}\{(u-\mu_*)^2/\sigma_*^2\}}{\sqrt{2\pi}}\,du$$

$$= \frac{\exp-\frac{1}{2}\{(\psi^2/\sigma_v^2)+(\delta' z/\sigma_u)^2-(\mu_*/\sigma_*)^2\}}{\sqrt{2\pi}(\sigma_u^2+\sigma_v^2)^{1/2}\{\Phi[\delta' z/\sigma_u]/\Phi[\mu_*/\sigma_*]\}}, \qquad \text{[A.9a]}$$

or, alternatively,

$$f_{\Psi}(\psi)=\frac{\exp-\frac{1}{2}\{(\psi-\delta' z)^2/(\sigma_v^2+\sigma_u^2)\}}{\sqrt{2\pi}(\sigma_u^2+\sigma_v^2)^{1/2}\{\Phi[\delta' z/\sigma_u]/\Phi[\mu_*/\sigma_*]\}}\,. \qquad \text{[A.9b]}$$

The density function for the cost value, vc_{ft}, in equation [A.1], is most conveniently given using the expression in equation [A.9b],

$$f_{VC_{ft}}(vc_{ft})=\frac{\exp-\frac{1}{2}\left\{\frac{[vc_{ft}-vc(x_{ft};\beta)-\delta' z]^2}{\sigma_v^2+\sigma_u^2}\right\}}{\sqrt{2\pi}(\sigma_u^2+\sigma_v^2)^{1/2}\{\Phi[d_{ft}]/\Phi[d_{ft}^*]\}}\,, \qquad \text{[A.10]}$$

where $d_{ft}=\delta' z_{ft}/\sigma_u$, $d_{ft}^*=\mu_{ft}^*/\sigma_*$ and $\mu_{ft}^*=[\sigma_v^2\delta' z_{ft}+\sigma_u^2(vc_{ft}-vc(x_{ft};\beta)]/(\sigma_u^2+\sigma_v^2)$.

Given that there are T_f observations obtained for the f^{th} firm, where $1 \leq T_f \leq T$, and $vc_f \equiv (vc_{f1}, vc_{f2}, ..., vc_{fT_f})'$ denotes the vector of the T_f cost values in equation [A.1], then the logarithm of the likelihood function for the sample observations, $vc \equiv (vc_1', vc_2', \ldots, vc_F')'$, is

$$\begin{aligned}L(\Theta^*;vc) = &-\frac{1}{2}\left(\sum_{f=1}^{F}T_f\right)\left\{\ln 2\pi+\ln(\sigma_u^2+\sigma_v^2)\right\}\\ &-\frac{1}{2}\sum_{f=1}^{F}\sum_{t=1}^{T_f}\left\{[vc_{ft}-vc(x_{ft};\beta)-\delta' z_{ft}]^2/(\sigma_u^2+\sigma_v^2)\right\}\\ &-\frac{1}{2}\sum_{f=1}^{F}\sum_{t=1}^{T_F}\left\{\ln\Phi[d_{ft}]-\ln\Phi[d_{ft}^*]\right\},\end{aligned} \qquad \text{[A.11]}$$

where $\Theta^* \equiv (\beta', \delta', \sigma_u^2, \sigma_v^2)'$.

Using the re-parameterization of the model suggested by Battese and Corra (1977), involving the parameters $\sigma^2 \equiv (\sigma_v^2 + \sigma_u^2)$ and $0 \leq \gamma \equiv \sigma_u^2/(\sigma_v^2 + \sigma_u^2) \leq 1$, the logarithm of the likelihood function can be expressed by

$$\begin{aligned}L(\Theta;vc) = &-\frac{1}{2}\left(\sum_{f=1}^{F}T_f\right)\left\{\ln 2\pi+\ln\sigma^2\right\}\\ &-\frac{1}{2}\sum_{f=1}^{F}\sum_{t=1}^{T_f}\left\{[vc_{ft}-vc(x_{ft};\beta)-\delta' z_{ft}]^2/\sigma^2\right\}\\ &-\frac{1}{2}\sum_{f=1}^{F}\sum_{t=1}^{T_f}\left\{\ln\Phi[d_{ft}]-\ln\Phi[d_{ft}^*]\right\},\end{aligned} \qquad \text{[A.12]}$$

where $d_{ft} = \delta' z_{ft} / (\gamma\sigma^2)^{1/2}$, [A.13]

$$d^*_{ft} = \mu^*_{ft} / [\gamma(1-\gamma)\sigma^2]^{1/2}, \quad \text{[A.14]}$$

$$\mu^*_{ft} = (1-\gamma)\delta' z_{ft} + \gamma(vc_{ft} - vc(x_{ft}; \beta)), \quad \text{[A.15]}$$

$$\sigma_* = [\gamma(1-\gamma)\sigma^2]^{1/2}, \quad \text{[A.16]}$$

and $\Theta \equiv (\beta', \delta', \sigma^2, \gamma)'$.

The log-likelihood function in equation [A.12] can be maximized with respect to each element of Θ to obtain maximum likelihood (ML) estimates of all parameters, β, δ, σ^2 and γ.

The computer program, FRONTIER Version 4.1, is used in this study to obtain the ML estimates for the parameters of the stochastic frontier cost model defined by equations [3]-[5] in the text. This program uses a three-step estimation procedure:

1. The first step involves calculation of the OLS estimators of β and σ^2.
2. In the second step, a grid search is conducted across the parameter space of γ, i.e., the log-likelihood function is evaluated for values of γ from 0.1 to 0.9 in increments of size 0.1. In these calculations, the β parameters (excepting β_0) are set to the OLS values, whit β_0 and σ^2 adjusted according to the corrected ordinary least squares formula presented in Coelli (1995). Any other parameters (δ-vector in our case) are set to zero during this grid search.
3. The final step uses the best estimates (that is, those corresponding to the largest log-likelihood value) from the second step as starting values in a Davidon-Fletcher-Powell (DFP) iterative maximization algorithm which obtain the final ML estimates when the likelihood function attains its global maximum.

Approximate standard errors of the ML estimators are then calculated by obtaining the square roots of the diagonal elements of the direction matrix from the final iteration of the DFP routine[82].

Once the ML estimates for the parameters of the stochastic frontier cost model have been obtained, predictions of the cost inefficiency for each producer, *f*, at each observation, *t*, have to be derived. We have estimates of $\psi_{ft} = v_{ft} + u_{ft}$ and we must extract the information that ψ_{ft} contains on the unobservable component u_{ft}. According to the original insight of Jondrow et al. (1982), a solution to the problem is obtained

[82] The direction matrix for the final iteration is usually a good approximation for the inverse of the Hessian of the log-likelihood function, unless the DFP routine terminates after only a few iterations.

from the *conditional distribution* of u_{ft} given ψ_{ft}, which incorporates whatever information ψ_{ft} contains concerning u_{ft}.

The conditional density function of u_{ft} given $\Psi_{ft} = \psi_{ft}$ is given by[83]

$$f_{U|\Psi=\psi}(u) = \frac{f_{\Psi,U}(\psi,u)}{f_{\Psi}(\psi)}, \qquad \text{[A.17]}$$

thus, using equations [A.6B] and [A.9b],

$$f_{U|\Psi=\psi}(u) = \frac{\exp - \frac{1}{2}\{(u-\mu_*)^2/\sigma_*^2\}}{\sqrt{2\pi}\sigma_*\Phi[\mu_*/\sigma_*]}. \qquad \text{[A.18]}$$

The overall cost *efficiency* of the f th firm at the tth observation, CE_{ft}, may be expressed as the ratio of stochastic frontier minimum cost (with $u_{ft} = 0$) to observed cost, which is equal to[84]

$$CE_{ft} = \frac{1}{\exp\{u_{ft}\}} = \exp\{-u_{ft}\}. \qquad \text{[A.19]}$$

This measure is bounded between zero ($u_{ft} \to \infty$) and one ($u_{ft} = 0$), and can be predicted in a similar way to that described for technical efficiency in the stochastic production frontier case analyzed by Battese and Coelli (1993). Using the conditional distribution of u_{ft} given ψ_{ft} defined by equation [A.17], the authors derive an expression for the conditional expectation of the technical efficiency for the f th firm at the tth observation, conditional upon the observed value of $\psi_{ft} = (v_{ft} - u_{ft})$. This expression, $E(\exp\{-u_f\}|\Psi_{ft} = \psi_{ft})$, is a generalization of the results presented in Jondrow et al. (1982) and Battese and Coelli (1988).

The prediction of the individual cost efficiencies relative to a stochastic cost frontier, i.e. expression [A.19], can be obtained by minor sign alterations of the technical efficiency point estimator in Battese and Coelli (1993). It is derived using the conditional density function of u_{ft} given $\Psi_{ft} = \psi_{ft}$ specified in equation [A.18] and is given by

$$C\hat{E}_{ft} = (\exp\{-u_{ft}\} | \Psi_{ft} = \psi_{ft}) = \left\{\frac{\Phi[(\mu^*_{ft}/\sigma_*)-\sigma_*]}{\Phi[\mu^*_{ft}/\sigma_*]}\right\}\exp\left\{-\mu^*_{ft} + \frac{1}{2}\sigma_*^2\right\} \qquad \text{[A.20]}$$

83 Again the subscripts, f and t, are omitted in the following expressions for convenience in the presentation.

84 Expression [A.19] is appropriate for CE_{ft} only if the general specification of the stochastic frontier cost model is given by equations [1]-[2] in the text.

where μ^*_{ft} and σ_* are defined by expressions [A.15] and [A.16].

Starting from the estimated stochastic cost frontier, FRONTIER Version 4.1 returns an estimate of cost *inefficiency* for each producer at each observation, i.e. a prediction of $CI_{ft} = \exp\{u_{ft}\}$. It measures the extent to which observed costs exceed the corresponding stochastic frontier values and it is then calculated as the inverse of CE_{ft} in equation [A.19], this last being predicted by applying the point estimator specified in equation [A.20].

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