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Regulatory contracts and cost efficiency: Stochastic frontier evidence from the Italian local public transport

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Abstract The main objective of this paper is to investigate the way subsidization mechanisms affect the cost efficiency of public transit systems, taking into account the role played by the environmental characteristics of each network. A cost frontier model is estimated for a seven-year panel of 44 Italian transit companies run under two different regulatory schemes (cost-plus or fixed-price), using the approach proposed by Kumbhakar et al. (1991), Huang and Liu (1994) and Battese and Coelli (1995). The main evidence is that, given network characteristics, transit operators with high-powered incentive contracts (fixed-price subsidies) exhibit lower distortions from the minimum costs. Environmental conditions (network speed levels) also have a significant impact on inefficiency differentials and influence the efficacy of incentive regulation. Overall, these results highlight a scope for transport policy to increase X-efficiency. Furthermore, they stress the importance of incentive theory and modern regulatory economics for the production analysis of regulated utilities.

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Introduction

A common feature of the regulatory framework of local public transport (LPT) in most countries is the provision for transfers from a regulation authority to the transit company. Since the latter generally face universal service obligations, commercial revenues are not high enough to cover operating costs. The payment of a subsidy is then required to ensure the balance of the budget. In most European countries, including Italy, this practice has led to a growing waste of public resources to cope with the consistent build up of deficits and the financial crisis faced by LPT firms. A better understanding of the sources of cost inefficiency in this industry may then be useful for reassessing traditional state intervention and designing new regulatory policies, in particular, with regard to contractual arrangements ruling the grant of subsidies.

So far few studies have explicitly analyzed the role of different subsidization mechanisms in explaining inefficiency differentials among LPT operators. The core of the present paper is to put forward information on the X-efficiency (Leibenstein 1966) of public transit systems in Italy, investigating the way regulatory contracts affect the distortions from the best-practice behavior of cost minimization. Since environmental characteristics of each network are likely to play an important role when dealing with cross section data to measure cost performance of transport companies, an attempt is made in the analysis to control for these effects using observable network variables. The results of this research are relevant for several reasons. First, they make a contribution to the debate as to whether the predictions from incentive theory (Laffont and Tirole 1993) help explain differences in productive efficiency among firms. Moreover, providing rigorous empirical evidence on the impact of different mechanisms for granting subsidies, our results have immediate policy implications for the ongoing regulatory reform in the Italian LPT sector, as will be explained in more detail below.

The analysis is based on a seven-year (1993–1999) balanced panel data of 44 Italian municipal companies managed under two regulatory schemes (i.e. cost-plus or fixed-price) and facing different levels of network commercial speed. The observed time period is particularly informative, as it encompasses both the years before and after the start-up of the reform. A stochastic frontier cost function is estimated applying the methodology developed by Kumbhakar et al. (1991), Huang and Liu (1994) and Battese and Coelli (1995). These authors propose an approach based on the parameterization of the mean of a truncated normal distribution, in which the inefficiency terms are a linear function of a set of explanatory variables including both firm-specific and time effects. In particular, the focus of the present study is on the way regulatory contracts, network characteristics, and their interaction affect X-inefficiency.

The rest of the paper is organized as follows. Section "Subsidies regulation and incentives to efficiency: The Italian framework" 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 "The econometric model", we develop the econometric model. Section "Specification of the stochastic frontier cost function", specifies the stochastic frontier cost function, while Section "Modelling inefficiency effects" deals with the modeling of cost inefficiency effects. The database is described in Section "Data description". Section "Empirical results" comments on the empirical results, discussing both the technology properties (Section "Technical characteristics") and the evidence on X-inefficiency (Section "Cost inefficiency and effects of regulatory schemes"). Section "Conclusion and policy implications" summarizes the major findings and provides some policy indications.

Subsidies regulation and incentives to efficiency: The Italian framework

During the first half of the nineties, many efforts were made with the aim of redressing the ruinous financial imbalance faced by the Italian LPT sector. Nevertheless, these interventions were only stopgap measures, that turned out to be inadequate to achieve the general goal of a structural readjustment of industry accounts. Only in the last seven years radical regulatory changes have been introduced in order to obtain the required improvements in terms of efficiency and effectiveness of the service and to reduce the waste of public funds spent on collective transports.[1] In 1995, Law 549 brought about 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 created specifically 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[2] and to stimulate the recovery of productive efficiency by transportation companies. Here it becomes necessary to eliminate the transfers from the central government and replace them with forms

[1] In 1995 about 71% of operating costs for the Italian bus-line companies came from public subsidies.

[2] *Decreto Legislativo* 422/1997 calls 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.

of taxation at regional level, in order to make binding any measures for an efficient use of public resources for local authorities. 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 company must obey as well as the reimbursement and the risk sharing scheme between the regulator and the operator.

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.[3] 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), the operator, as it is not residual claimant for effort, has no incentives to produce efficiently. Since 1996, some municipalities have introduced alternative reimbursement mechanisms that, even if not yet formalized within a proper service contract, have virtually overcome the ex-post balancing of accounts: the *gross cost schemes*, under which the industrial risk is entirely borne by the firm while the commercial risk is borne by the local authority, and the *net cost schemes*, that provide for the assumption of both types of risk by the operator. These two types of contractual arrangements are traceable to what the theory of incentives in regulation calls 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 companies subjected to fixed-price mechanisms are assumed to face high-powered incentives towards a cost minimizing behavior.[4]

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.[5] The complex problem of designing an optimal contract is beyond the scope of our study,[6] since only fixed-price or cost-plus schemes are carried out 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 whether transit companies run under fixed-price regimes are more cost efficient due to the fact that they face stronger incentives to increase managerial effort. If this 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 us to assess whether 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.

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

[3] The operator Bolzano SASA is the only exception, as it has already been subjected to a standard-cost regime since 1989.

[4] Actually, with the second-type of fixed-price scheme, the so-called net cost contract, the regulated company is responsible for both insufficient revenues and cost overruns. Therefore, it has an incentive to increase traffic proceeds besides reducing operating costs, for instance by raising the quality of the service or controlling more severely the tariff evasion. However, since the focus of this study is on cost efficiency of the supply, we will not deal with issues concerning informational asymmetries between regulator and LPT operator on the demand side and related incentive problems.

[5] See Laffont and Tirole (1993) for a complete description of this problem.

[6] To this regard, see Wunsch (1994b), Dalen and Gomez-Lobo (1996, 1997) and Gagnepain and Ivaldi (2002b).

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:

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

$$\text{with } \psi_{ft} = v_{ft} + u_{ft}, \quad (2)$$

where VC denotes the 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 ($t = 1, \ldots, T$).

As usual in 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 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 fth firm at the tth 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, the cost frontier is attained.

Specification of the stochastic frontier cost function

In order 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 suggested in the literature (e.g., Windle 1988; Levaggi 1994; Fabbri 1998; Fraquelli et al. 2004), the rolling stock should be considered as a fixed factor in the short-run. The model includes: a scalar output (Y); the price of three variable factors, i.e. labor (L), fuel (F), materials and services (MS); a quasi-fixed input (K); three network variables, i.e. the average commercial speed (SP), a dummy for intercity companies ($DINTC$) and for "mixed" operators, which supply in combination urban and intercity public transport ($DMIX$); a time trend variable (τ).

As for the choice of the output indicator, we decided to ignore the demand side in the specification of the cost model for a twofold reason. First, even if we are conscious that transit demand is important for the long-run equilibrium of the industry, however, as already suggested by numerous authors, passengers are not under the direct control of the firms in the short run (Small 1990; Berechman 1993). Moreover, till now in Italy LPT operators have been involved in providing a specific production capacity determined by a certain number of seats available and a certain mileage performed by vehicles, so that the capacity (intermediate output) is a typical cost driver, while passenger-trips (final output) represent mostly a revenue driver. According to this reasoning, we specified the output as the number of seat-kilometers offered.[7] This indicator, obtained by multiplying the average load capacity of vehicles by the total traveled kilometers, is particularly suited to our firm sample, which includes both urban and intercity services.[8]

The rolling stock plays the role of quasi-fixed input in our variable cost model. Following a specification extensively used in the transportation literature

[7] The majority of frontier analyses dealing with European LPT data have specified the output measure in terms of seat-kilometers. Applications of parametric methodologies include, among the others, Delhausse et al. (1992), Filippini et al. (1992), Thiry and Tulkens (1992) and Fazioli et al. (1993), while the contributions by Gathon (1989), Tulkens (1993), Wunsch (1994a) and Tulkens and Vanden Eeckaut (1995) are examples of studies based on non-parametric methodologies (DEA and FDH).

[8] Intercity firms generally perform a higher number of kilometers than the urban units, as they cover a larger network; on the other hand, a urban company reasonably offers a higher number of places (buses usually are larger and also their number is higher, because there is a more intensive demand to satisfy). Compared to the alternative supply indicator proposed by the LPT literature, i.e. the vehicle-kilometers (De Borger et al. 2002), the output measure we adopt allows us to weigh the specific characteristics of urban and intercity systems, as it takes also into account the average load capacity of the fleet.

(McCarthy 2001), capital was defined as the total number of vehicles owned by each company.

Prices of variable factors were computed using 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 of diesel oil consumed.[9] Expenses for materials and services represent a residual input category and mostly refer to costs of spares and of repair and maintenance services bought outside. Since only very disintegrated data would allow to separate the different measures of inputs and the related prices, it is reasonable to assume that these expenses strictly depend on the actual exploitation of the rolling stock; consequently, we adopted the solution proposed by Gagnepain and Ivaldi (2002b) and derived an average price for this composite input (P_{MS}) by dividing costs of materials and services by the average number of vehicles used in the year.[10]

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 previous works on the industry (e.g., Windle 1988; Levaggi 1994; Wunsch 1996; Gagnepain 1998; Fraquelli et al. 2004). 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. Moreover, since our sample embraces LPT companies providing urban, intercity, or both types of service, we added two *service-specific* dummies for intercity and mixed firms, so as to consider possible shifts of frontier cost levels due to different network configurations. Dummy *DINTC* assumes value 1 for intercity companies and 0 otherwise, similarly *DMIX* is equal to 1 in the case of mixed operators and 0 for specialized networks.

Finally, we included a time trend in the model, 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 possible scientific or organizational progress should not be negligible. Assuming the other things are unchanged, costs are then expected to diminish over time. Summary statistics (mean, variability index, minimum and maximum) on the variables used in the empirical application are provided in Table 1.[11]

A translog functional form is chosen for the analysis.[12] The stochastic frontier cost model (1)–(2) is then defined by the equation:

$$
\begin{aligned}
\ln\left(\frac{VC_{ft}}{P_{Fft}}\right)
&= \beta_0 + \beta_y \ln Y_{ft} + \beta_k \ln K_{ft} + \sum_i \beta_i \ln\left(\frac{P_{ift}}{P_{Fft}}\right) \\
&\quad + \beta_{SP} \ln SP_{ft} + \sum_i \beta_{iy} \ln\left(\frac{P_{ift}}{P_{Fft}}\right) \ln Y_{ft} \\
&\quad + \sum_i \beta_{ik} \ln\left(\frac{P_{ift}}{P_{Fft}}\right) \ln K \\
&\quad + \sum_i \beta_{iSP} \ln\left(\frac{P_{ift}}{P_{Fft}}\right) \ln SP_{ft} + \beta_{yk} \ln Y_{ft} \ln K_{ft} \\
&\quad + \beta_{ySP} \ln Y_{ft} \ln SP_{ft} + \beta_{kSP} \ln K_{ft} \ln SP_{ft} \\
&\quad + \frac{1}{2}\beta_{yy}(\ln Y_{ft})^2 + \frac{1}{2}\beta_{kk}(\ln K_{ft})^2 \\
&\quad + \frac{1}{2}\beta_{SPSP}(\ln SP_{ft})^2 + \frac{1}{2}\sum_i \sum_j \beta_{ij} \ln\left(\frac{P_{ift}}{P_{Fft}}\right) \\
&\quad \times \ln\left(\frac{P_{jft}}{P_{Fft}}\right) + \beta_{INTC} DINTC_{ft} + \beta_{MIX} DMIX_{ft} \\
&\quad + \beta_\tau \tau_{ft} + v_{ft} + u_{ft} \quad i, j \in \{L, MS\},
\end{aligned}
\tag{3}
$$

where the normalization of the monetary variables, VC, P_L and P_{MS}, with respect to the price of fuel,

[9] For a few firms which utilize tramways, trolley-lines or railways and consume electricity, kilowatt-hours were transformed in equivalent-liters of diesel oil.

[10] This measure was obtained by multiplying the total number of vehicles in the rolling stock by the average annual rate of fleet utilization. The latter was provided by each company through a questionnaire (see Section "Data description").

[11] Table 1 shows large variations in the values of variable costs, output and capital, which partially reflect the presence of different operating conditions for urban, intercity and mixed firms.

[12] 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 underlying 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.

Table 1 Summary statistics for the variables of the stochastic frontier cost function

	mean	var. index[a]	min	max
VC[b] (millions Lire)	68,606	1.24	654	456,049
Y[c] (millions)	997	1.14	6	6,554
K[d]	270	0.99	5	1,581
P_L (millions Lire / worker)	73.69	0.09	55.38	90.55
P_F (Lire / liter of diesel oil)	1,127	0.12	662	1,752
P_{MS} (millions Lire / vehicle)	44.21	0.37	13.74	111.22
SP (Kms/h)	23.04	0.38	13.00	47.00

[a] The variability index (*var. index*) is the ratio of the standard deviation to the mean
[b] Variable operating cost (VC): sum of labor, fuel, and materials and services expenses
[c] Output (Y): seat-kilometers offered in each year
[d] Capital (K): total number of vehicles in the rolling stock

P_F, is made to ensure the linear homogeneity of the cost function in input prices.[13]

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.

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 et al. (1977) and Meeusen and van den Broeck (1977).[14] In particular, during the last decades some authors have proposed different methods to investigate the *determinants* of inefficiency differentials among firms. This issue was initially tackled with a two-step approach, by which inefficiency and exogenous effects were identified sequentially (e.g., Pitt and Lee 1981; Kalirajan 1981; Kalirajan and Shand 1989).[15] Successively, Kumbhakar et al. (1991), Reifschneider and Stevenson (1991) and Huang and Liu (1994) specified stochastic frontier models in which the inefficiency effects were defined as explicit functions of some variables involving firm characteristics (e.g., the degree of competitive pressure, input quality indicators, various managerial factors) and all parameters were estimated in a single-stage maximum likelihood (ML) procedure. Starting from Kumbhakar (1990), in which productive inefficiency is allowed to vary over time, Battese and Coelli (1995) adapted the one-step approach to accommodate panel data, which permits to include both firm-specific and time effects in the model adopted to explain inefficiencies. From a methodological perspective, the studies by Kumbhakar et al. (1991), Huang and Liu (1994) and Battese and Coelli (1995) have in common the feature of parameterizing the mean of a truncated normal distribution as a way to analyze the exogenous influence on inefficiency.[16] Thus, as suggested by Wang and Schmidt (2002), we can call them Kumbhakar–Ghosh–McGuckin–Huang–Liu–Battese–Coelli model (KGMHLBC hereafter).

13 Symmetry property ($\beta_{ij} = \beta_{ji}$ 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.

14 An up-to-date and detailed review of the literature on stochastic frontier modeling and efficiency measurement is provided in Kumbhakar and Lovell (2000).

15 The drawbacks of this procedure are highlighted in Coelli et al. (1998, Chapter 9) and Kumbhakar and Lovell (2000, Chapter 7).

16 Another approach to investigate the exogenous determinants of inefficiency, which complements the modeling strategy of KGMHLBC, is that of Caudill and Ford (1993), Caudill et al. (1995) and Hadri (1999). These contributions (CFCFGH model) address the problem of heteroscedasticity of the inefficiency effects by parameterizing the variance of a truncated normal distribution. For a discussion of the properties of KGMHLBC and CFCFGH models see Kumbhakar and Lovell (2000).

Using the «FRONTIER Version 4.1» computer program by Coelli (1996),[17] this study applies the KGMHL BC methodology for analysing the effects of regulatory constraints on the cost efficiency of public transit systems in Italy, taking into account the role played by the environmental characteristics of each network. The objective, in particular, is to investigate whether the predictions from the theory of incentives in regulation (Laffont and Tirole, 1993) help to explain differences in productive efficiency, i.e. *do high powered incentive regulatory schemes increase efficiency as compared to low powered schemes?* In the recent survey of the literature on production and cost frontiers for public transit systems by De Borger et al. (2002), the authors argue that most studies find that productive inefficiencies are widespread in the sector and that technical inefficiency represent the main source of poor performance, rather than congestion or scale inefficiency. They also note that frontier evidence shows a relevant influence of both the regulatory environment and the network characteristics on X-efficiency and productivity, highlighting a need for comparative international research to provide more details on the relative performance of firms run under different regulation regimes. Indeed, as Dalen and Gomez-Lobo (2003) underline, few works have explicitly investigated the role of subsidization practices in explaining variations of measured efficiency among transit companies and a still more restricted number of them has empirically analyzed how firms respond to changes in regulatory incentives.

As for Europe, the econometric studies on this subject include Dalen and Gomez-Lobo (1996, 1997) for public transit system in Norway and Gagnepain and Ivaldi (2002a, b) for the France. They specify a Cobb–Douglas technology and use a structural approach to recover the underlying cost efficiencies associated with different regulatory contracts. However, due to the simplified functional form and the structural nature of the estimation strategy, these studies appear rather restrictive and may thus introduce unnecessary biases. To avoid these problems, here we decided to adopt a reduced form approach based on the estimation of a flexible translog cost frontier model.[18] In the remainder of the section, the principal features of the theoretical framework we will refer to in the empirical analysis of X-efficiency are briefly resumed, focusing the attention on the role of *informational asymmetries* and incentives in the regulator-firm relationship stressed by the new regulatory economics. Then we will describe the variables we suppose capture these effects within the KGMHLBC model. The discussion concerning the estimation procedure is postponed to the following section.

Taking cue from the findings of the theory of incentives in regulation (Laffont and Tirole 1993) and their applications to the LPT sector (Dalen and Gomez-Lobo 1996, 1997; Gagnepain and Ivaldi, 2002a, b), global cost inefficiency, or X-inefficiency, can be interpreted as the combined result of the presence of exogenous technical inefficiency—supposed to imply a fall in the productivity of labor input[19] —and of the cost-reducing activity exerted by managers to counterbalance the negative effect of the intrinsic lack of labor productivity.[20] It is reasonable to suppose that labor inefficiency is perfectly known by the firm and not known or imperfectly known by the regulatory authority (*adverse selection* phenomenon), as the latter does not take

[17] This software allows to specify the stochastic frontier model in terms of a dual cost relationship instead of a production function.

[18] A reduced form approach is used also in Dalen and Gomez-Lobo (2003) to study the effects of the introduction of more high powered incentive schemes based on a yardstick type of regulation.

[19] Labor input, which represents the highest share of total operating costs (seventy percent on average in Italy), is the major source of informational asymmetries. This assumption is based on the view that bus drivers play a decisive and acute role in operating the network, especially with respect to the flexibility and punctuality of operations in peak period. Labor inefficiency is assumed to be "exogenous" (or "intrinsic") in the sense that it is given and cannot be changed by managers in the short run, as it depends on factors such as the geographical and historical characteristics of the network, the structure of the labor force or the ability level of drivers. For more details on this issue (see Ivaldi 2000, pp. 740–75).

[20] The operator may spend time and effort on monitoring workers, for instance providing drivers with training programs, solving potential conflicts among them, avoiding strikes, etc. From a microeconomic point of view, it is worthwhile to underline that, unless the effort level of managers fully offsets intrinsic labor inefficiency, the firm minimizes costs by taking into account a higher labor price. That is, marginal rate of technical substitution of any pair of inputs containing labor is higher than the observed price ratio of inputs, which may lead to cost inefficiencies.

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 (*moral hazard* phenomenon).[21] Since there is no clear evidence on the motivation of a manager to work as hard as he could, it is precisely in this context that the type of regulatory scheme and related incentives faced by the firm during the production process play a role in reducing overall cost inefficiency.

Within the described framework, the KGMHLBC stochastic frontier model is applied to the analysis of cost inefficiency (from Eqs. (1)–(2): $\exp\{u_{ft}\}$) of public transit systems in Italy. The emphasis is put on the impact of the different subsidization mechanisms that companies have to face, also taking into account the specific characteristics of each network in a way which will be specified later. The specific formulation of KGMHLBC model proposed by Battese and Coelli (1995) defines the X-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[22] and may also include some variables involved in the cost function (provided the inefficiency effects are stochastic) and/or interactions between these latter and firm-specific factors. The X-inefficiency effect incorporated in the composed error term, ψ_{ft}, of the general stochastic frontier model (1)–(2) could be specified by the equation:

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

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.

Since our objective is to verify whether the causes of X-inefficiency affecting the Italian public transit systems should be searched for in the system of incentives generated by the regulatory environment, we first introduce a regulation dummy, R_{ft}, as determinant of u_{ft}. As previously mentioned, 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 as to allow ex-post budgets to be balanced,[23] and fixed-price schemes, where the transit operator obtains a transfer defined ex-ante in order to finance an expected operating deficit.[24] 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, cost inefficiencies are expected to be significantly lower under fixed-price regulation, because 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 such as the geographical and historical characteristics influencing the structure and the operability of the network,

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

[22] 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.

[23] In Section "Subsidies regulation and incentives to efficiency: The Italian framework" we refer to this reimbursement rule as the "management contract", under which the regulated firm does not bear any risk.

[24] This type of reimbursement rule specifically refers to the subsidization scheme we have previously called "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 bears 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.

the ability level of drivers, the public policy for local mobility, etc. These elements jointly contribute to determine what we have called above intrinsic labor inefficiency, or exogenous technical inefficiency. Even if we do not obtain here a specific estimate for this unobservable component of the global cost inefficiency,[25] an attempt is made to include in the inefficiency model (4) a variable strictly related to the above factors (hereafter "network characteristics"), likely to influence the exogenous technical inefficiency and then the level of the overall cost distortion. To this end we introduce as a proxy for network 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.[26] Since a higher value for this variable is supposed to reflect better operating conditions,[27] thus reducing the intrinsic inefficiency level, we expect to find a negative sign for the coefficient associated with $\ln SP_{ft}$, δ_{SP}, in model (4). Furthermore, in order to take into account the possibility that when the exogenous inefficiency of a network is too high the cost-reducing activity exerted by managers could have a little weigh in determining global cost efficiency and the role of contractual arrangements becomes then modest,[28] the impact of 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 $(R_{ft} \ln SP_{ft})$.[29] 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.[30]

As the Battese and Coelli formulation of KGMHLBC approach enables us to include both firm-specific and time effects in the specification of inefficiency model, we also incorporate in the Eq. (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 that began 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 as to allow the dynamics of cost inefficiencies throughout the analyzed period to vary with the regulatory pattern. We denote this variable with $(R_{ft}\tau_{ft})$, 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:

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

Equation (5) indicates that stochastic X-inefficiency effects are linearly related to the regulatory scheme and commercial speed of the transit companies, the period of observation, and the interactions of speed and time with regulation, with an intercept parameter included in the model.

[25] We remark that the intrinsic labor inefficiency represents an *adverse selection* variable which reflects private information on the firm's technology that is not known (or imperfectly known) by the regulator and the econometricians.

[26] Transformation in logarithms is maintained for homogeneity with the Eq. (3). In both cases the logarithm specification enables the interpretation of the partial derivatives of the dependent variables, $\ln VC_{ft}$ and u_{ft}, computed with respect to $\ln SP_{ft}$ in terms of elasticities.

[27] 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.

[28] In fact, evidence supporting this conjecture has been found in the study of Gagnepain and Ivaldi (2002a, b). The authors show that fixed-price mechanisms generally provide more incentive for efficiency, but contractual arrangements do not appear to be very relevant for the firm's performance when the operators are characterized by a fairly high intrinsic inefficiency.

[29] A similar specification is adopted in Bhattacharyya et al. (1995) to analyze the impact of ownership structure on cost efficiency of public transit systems in India. The authors include in the inefficiency model ownership dummies as well as their interactions with firm-specific characteristics affecting inefficiency, i.e. the rates of breakdown and vehicle utilization.

[30] At the same time, a negative sign for δ_{RSP} would mean that the effects on cost efficiency due to a gain in the average commercial speed are strengthened in the presence of fixed-price schemes, because of the higher cost-reducing effort provided by managers under this type of regulation.

Distributional assumptions and estimation procedure

The final stochastic frontier model to be estimated is specified in Eq. (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 $\sim$ $i.i.d.$ $N(0, \sigma_v^2)$, 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, so that u_{ft} arises from the truncation (at zero) of the normal distribution with mean $\delta' z_{ft}$ and variance σ_u^2, $N(\delta' z_{ft}, \sigma_u^2)$.[31] This assumption allows individual cost inefficiencies to depend on firm- and time-specific exogenous observable factors, z_{ft}.

The ML method is employed for the simultaneous estimation of parameters of the stochastic frontier [3] and the model for the cost inefficiency effects [5]. 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 $\gamma \equiv \sigma_u^2/(\sigma_v^2 + \sigma_u^2)$.[32] The parameter γ must lie between 0 and 1 and provides a useful indication of the relative contributions of u_{ft} and v_{ft} to ψ_{ft}. As $\gamma \to 0$ the symmetric noise component, v_{ft}, dominates the one-sided cost inefficiency term, u_{ft}, in determining the variation of global residual, ψ_{ft}. The inverse occurs as $\gamma \to 1$. In the former case we are back to a traditional average cost function model with no stochastic inefficiency, whereas in the latter case we are back to a deterministic cost frontier model with no random noise included.[33]

After obtaining parameter estimates, we consider the estimation of u_{ft}. When the model in Eq. (4) is assumed, the overall cost inefficiency of production for the fth firm at the tth observation is defined by the expression:

$$CI_{ft} = \exp\{u_{ft}\} = \exp\{\delta' z_{ft} + w_{ft}\}, \qquad (6)$$

which takes a value between one (when $u_{ft} = 0$) and infinity (when $u_{ft} \to \infty$). The prediction of the X-inefficiencies, $\widehat{CI}_{ft}$, is based on conditional expectations which generalize the estimators computed in Jondrow et al. (1982) and Battese and Coelli (1988).[34]

Data description

The study uses a seven-year balanced panel data of 44 companies operating in the Italian LPT sector from 1993 to 1999, for a total of 308 observations. All firms are members of ASSTRA (Rome), a nationwide trade organization which associates the Italian publicly-owned LPT operators.[35]

The sample composition by type of service is the following: 17 firms mostly operate in the urban context, 10 mainly provide intercity service, and the remaining 17 have activities in both compartments. As far as the distribution by geographical area is concerned, the sample is fairly balanced: 25 operators are located in the North regions and 19 in the Center-South regions (in particular, 10 in the Center and 9 in the South). The prevalence in the sample of companies providing only bus service (38 units) compared to the multi-modal firms (6 units, supplying also tramways and railways) reflects the modality composition at national level,

[31] This can also be written as $u_{ft} \sim N^+(\delta' z_{ft}, \sigma_u^2)$, where $N^+(\delta' z_{ft}, \sigma_u^2)$ indicates a truncated-normal distribution with mode $\delta' z_{ft}$ and spread parameter σ_u^2. With regards to the concepts of half-normal and truncated-normal distribution see Kumbhakar and Lowell (2000), pp. 74–86.

[32] The log-likelihood function for a *production* frontier model is presented in the Appendix of Battese and Coelli (1993). The *cost* frontier version requires a few sign changes and is derived in Piacenza (2002).

[33] The term *deterministic* is used because in this type of frontier model the observed cost, VC_{ft}, is bounded below

Footnote 33 continued
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.

[34] This result is also provided in the Appendix of Battese and Coelli (1993). It is worthwhile pointing out 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 $t \neq t'$. Thus it follows that the same ordering of firms in terms of cost inefficiency of production does not apply to all time periods.

[35] In 2000 the members of ASSTRA came to around 165, equal to 90% of the urban operators and to 50% of the intercity operators in Italy (Boitani and Cambini 2002). The sample we use may then be considered to be sufficiently representative of the Italian LPT industry.

Table 2 Number of operators run under each type of regulatory scheme

	1993	1994	1995	1996	1997	1998	1999	All years
Cost-plus contracts	44	44	44	27	27	27	19	232
Fixed-price contracts	0	0	0	17	17	17	25	76

where the road mode of transportation represents about 80% of LPT services in terms of seat-kilometers.[36] As for firm size, measured in terms of the average number of employed workers, the sample includes 17 large-sized firms (more than 550 workers), 19 medium-sized units (151–550 workers), and 8 small operators (less than 150 workers). Finally, as far as the subsidization mechanisms are concerned, 25% of observations (76 cases) relate to fixed-price regulatory schemes, while 75% (232 cases) refer to transit systems under cost-plus reimbursement rules. The sample structure by regulation regime in each year is shown in Table 2.

For the panel construction we resorted to two different informational sources. The starting database gathers information extracted from ASSTRA annual reports concerning the years indicated above. From these reports we were able to derive the 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. The data was appropriately integrated by further information on cost, technical-environmental factors and type of regulatory scheme obtained through questionnaires sent to the companies. This additional investigation enabled the cost to be split by productive factors other than labor, such as fuel, materials and services, and capital. Moreover, we retrieved relevant technical information (average load capacity of each vehicle, annual rate of fleet utilization, network commercial speed), in order to complement the data extracted from the ASSTRA annual reports. To analyze 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, we also included in the questionnaire a question on the reimbursement mechanism adopted by the competent local authority (Region, Province or Town Council).[37] This rich source is probably unique in Italy as a tool of comparing regulatory systems to each other and over time.

Empirical results

ML estimates for parameters of the model defined in Eqs. (3)–(5) are given in Tables 3 and 4. In particular, Table 3 reports the estimated coefficients, β, for the stochastic frontier cost function (3), while Table 4 presents the estimates of the inefficiency-related coefficients, δ, for the model [5] and the two variance parameters, γ and σ^2.

In order to check if the translogarithmic functional form gives an adequate representation of the cost structure of our sample of LPT firms, we run generalized likelihood ratio (LR) tests on the technology restrictions implied by a Cobb-Douglas specification ($\beta_{ij} = \beta_{iy} = \beta_{ik} = \beta_{iSP} = \beta_{yy} = \beta_{kk} = \beta_{SPSP} = \beta_{yk} = \beta_{ySP} = \beta_{kSP} = 0$) and by homotheticity ($\beta_{iy} = \beta_{ik} = \beta_{iSP} = 0$). The test statistic is asymptotically distributed as a chi-square (χ^2) random variable with degrees of freedom equal to the number of restrictions involved. LR tests results reported in Table 5 lead to reject both restrictions and to retain the general model (3).[38]

Most of β coefficients are larger than their estimated standard errors and are statistically significant. The signs of the first-order 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

[36] Source: Ministry of Transports and Navigation (1997).

[37] 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.

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

Table 3 Maximum-likelihood estimates for parameters of the stochastic frontier cost function (3)[a]

Regressors	Parameters	Estimates	Standard errors
Constant	β_0	17.797***	0.033
ln Y	β_y	0.518***	0.048
ln K	β_k	0.054	0.050
ln P_L	β_L	0.633***	0.063
ln P_{MS}	β_{MS}	0.119***	0.029
ln SP	β_{SP}	−0.264***	0.048
ln P_L ln Y	β_{Ly}	0.320	0.269
ln P_{MS} ln Y	β_{MSy}	−0.211	0.130
ln P_L ln K	β_{Lk}	−0.197	0.298
ln P_{MS} ln K	β_{MSk}	0.122	0.134
ln Y^2	β_{yy}	0.448****	0.139
ln K^2	β_{kk}	0.260	0.173
ln Y ln K	β_{yk}	−0.355**	0.150
ln P_L ln P_{MS}	β_{LMS}	−0.516**	0.206
ln P_L^2	β_{LL}	1.451****	0.510
ln P_{MS}^2	β_{MSMS}	0.291***	0.102
ln Y ln SP	β_{ySP}	−0.325***	0.095
ln K ln SP	β_{kSP}	0.316***	0.114
ln P_L ln SP	β_{LSP}	−0.488***	0.137
ln P_{MS} ln SP	β_{MSP}	0.145*	0.079
ln SP^2	β_{SPSP}	−0.270*	0.150
DINTC	β_{INTC}	−0.054**	0.023
DMIX	β_{MIX}	−0.075**	0.034
τ	β_τ	−0.014*	0.003

[a] All the independent variables excepting time have been normalized to their sample mean before the log-transformation
*Statistically significant at the 10% level
**Statistically significant at the 5% level
***Statistically significant at the 1% level

Table 4 Maximum-likelihood estimates for parameters of the stochastic cost inefficiency model (5)[a]

Regressors	Parameters	Estimates	Standard errors
Constant	δ_0	−0.604**	0.243
R	δ_R	−1.521***	0.478
ln SP	δ_{SP}	−0.468***	0.149
τ	δ_τ	0.036**	0.016
R ln SP	δ_{RSP}	−0.896***	0.275
R^τ	$\delta_{R\tau}$	−0.014**	0.006
Sigma-squared	σ^2	0.083****	0.023
Gamma	γ	0.969****	0.014

[a] All the independent variables excepting time have been normalized to their sample mean before the log-transformation
**Statistically significant at the 5% level
***Statistically significant at the 1% level

Table 5 Likelihood-ratio tests for parameters of the stochastic frontier cost function (3)

Null hypothesis	Log-likelihood	χ^2-statistic	Decision
H_0: Cobb-Douglas specification ($\beta_{ij} = \beta_{iy} = \beta_{ik} = \beta_{iSP} = \beta_{yy} = \beta_{kk} = \beta_{SPSP} = \beta_{yk} = \beta_{ySP} = \beta_{kSP} = 0$)	217.641	85.281	Reject H_0
H_0: Homotheticity ($\beta_{iy} = \beta_{ik} = \beta_{iSP} = 0$)	230.843	29.438	Reject H_0

microeconomic theory.[39] An intense debate has arisen in literature on this problem. According to Filippini (1996), the positive sign of K is due to a problem of multicolinearity in cases where a positive correlation between dependent variable and capital indicator exists. The alternative argument put forward by Caves et al.(1985) in a study on the U.S. railroads is that the positive sign of K reflects an industry that does not minimize costs in the long term and therefore employs too much capital in the production process. This interpretation has been afterwards extended to public transit systems in U.S. (Windle 1988; McCarthy 2001), Belgium (Tulkens et al. 1988), and Italy (Levaggi 1994). In these studies the authors argue that the inefficient use of capital could derive from the generous government programs of subsiding investments; such a way of providing capital grants distorts the input allocation: by effectively reducing the cost of purchasing additional vehicles, these capital subsidies provide economic incentives for overcapitalization, as LPT operators substitute the relatively less expensive buses for other inputs in the production of transit services. The presence of considerable excess capacity has important implications for the regulation design concerning future investments. As remarked by McCarthy (2001), LPT companies can be driven to move towards the optimal rolling stock size by temporarily cutting off the financial aids and forcing them to an attrition policy: as vehicles are retired from service, operators are constrained to more efficiently utilize the remaining fleet instead of purchasing new capital. This would yield cost savings to the firms, primarily by reducing extra maintenance costs due to unused capacity.

The δ coefficients associated with the explanatory variables in the inefficiency model (5) are of particular interest to this study. The hypotheses that cost inefficiency effects are absent or that they have simpler distributions have been statistically tested using the generalized LR test. 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 the null hypothesis is true, the LR statistic has asymptotic distribution which is a mixture of χ^2 distributions whose critical values are obtained from Table 1 in Kodde and Palm (1986).[40] 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% level of significance. 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}, $(R_{ft} \ln SP_{ft})$ and $(R_{ft}\tau_{ft})$, are included in the cost function.[41] Also this hypothesis is rejected at the 1% level. Similarly, the null hypotheses that the $u_{ft}s$ are altogether unrelated to the z-variables, that they are not a linear function of subsidization mechanisms, average network speed, year of observation, and interaction of regulation with speed and time, and that they do not include an intercept parameter are all also rejected at the 1% level.

The estimates for δ parameters of the restricted model (5) are presented in Table 4. The coefficients have the expected sign and are all statistically significant. The values of parameters $\sigma^2 \equiv (\sigma_v^2 + \sigma_u^2)$ 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 estimate for γ is 0.969 with asymptotic standard error of 0.014. The results is consistent with the conclusion that the true γ-value is accepted to be greater than zero (in the LR tests above), showing that the vast majority of residual variation is due to the inefficiency effects and that a traditional average response function would not adequately represent the data. However, we also see that the γ estimate is significantly less than one, to indicate that our stochastic frontier model (3)–(5) may be significantly different from a deterministic

[39] A variable cost function should also satisfy the property that is non-increasing with respect to capital stock (see Cornes 1992, p. 106). It is worthwhile noticing that the wrong sign estimated for the capital coefficient seems to be a general problem characterizing the use of a variable cost function model, not only in the studies on transportation industry.

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

[41] Note that the parameters δ_0, δ_{SP} and δ_τ cannot be identified if there are no random inefficiency effects in the model ($\gamma = 0$), as 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.

frontier specification, in which there are no random errors, v_{ft}, in the cost function.[42]

In the following Section "Technical characteristics" we will take a brief look at the frontier cost elasticities and at the technological properties of the estimated translog model.[43] We postpone to Section "Cost inefficiency and effects of regulatory schemes" the discussion concerning detected X-inefficiencies and the effects of regulatory schemes, which are our primary interest in this study.

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 Table 3 can be interpreted as frontier cost elasticities for the *average operator* of the industry.[44] 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}, type of service, β_{INTC} and β_{MIX}, and time, β_τ. These have been utilized to infer the characteristics of technology (evaluated at the sample mean) presented in Table 6, where the separated effects on frontier costs attributable to short-run (SRS) and long-run returns to scale (LRS), commercial speed improvements, shift of firm's production from urban to intercity and mixed services, and Hicks-neutral technological change are highlighted.

The analysis reveals the presence of short-run and long-run scale economies. Indeed, asymptotic t-tests lead to the acceptance of both hypotheses that SRS and LRS are significantly greater than one. The estimated SRS, 1.93, show that, given the endowment of quasi-fixed input, a more than proportional 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. As far as LRS are concerned,[45] the estimate, 1.83, implies a sub-optimal scale with respect to the long-run equilibrium. On the whole, these results highlight 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.

The estimated frontier cost elasticity with respect to the average speed of the network in Table 6 ($\beta_{SP} = -0.26$) 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% brings about the reduction in the level of operating costs for the average firm by 2.6%. This result underlines the importance of appropriate public policies concerning local traffic regulation.

Third column of Table 6 reports service-specific cost elasticities associated with the dummies for the intercity (DINTC) and mixed (DMIX) activity.[46] Both estimates have a negative sign and are statisti-

[42] Note that the very high estimate obtained for γ may appear a surprising result for a public transit cost function, where one would normally expect data noise to play a larger role, given that bus companies are likely to have different networks and a variety of unobserved environmental factors which might affect their cost. However, this evidence is rather common in the applications of KGMHLBC model (see the stochastic frontier studies on agricultural sector and telecommunications quoted in Coelli et al. (1998), where the value of γ is always very close to one) and it is consistent with the findings on public transit by Dalen and Gomez-Lobo (2003), who have obtained estimates of γ ranging between 0.85 and 0.93. From an economic perspective, it is worthwhile to underline that the major role played by the inefficiency term can also be due to the specific context of regulated public utilities; as argued by Kumbhakar (2000), they typically operate in noncompetitive environments, where the possibility of X-inefficiency is wider compared with non regulated industries, in which the firms are less constrained in their ability to freely adjust their use of inputs. However, the fact that γ is close to one might be also because of a shortcoming of our model, as it does not account for unobserved heterogeneity among firms in the specification of the cost frontier, which might inflate the variance of the error term, in particular that of the modeled inefficiency component. I thank an anonymous referee for having raised such a critical issue.

[43] These aspects are also analyzed and discussed in more details in Cambinie and Filippini (2003) and Fraquelli et al. (2004).

[44] The *average operator* (the point of normalization) corresponds to a hypothetical firm operating at an average level of production, using an average stock of quasi-fixed input, and

Footnote 44 continued
facing average variable input prices and average commercial speed over the sample.

[45] We can evaluate the long-run returns to scale by applying the algorithm first suggested by Caves et al. (1981) and indicated in square brackets in Table 6.

[46] Service-specific cost elasticities represent the percentage effect on variable costs due to the shift of the firm's production from urban to intercity or mixed service. The computation follows Halvorsen and Palmquist (1980).

Table 6 Technology characteristics evaluated at the mean of the data (average firm)[a]

Returns to scale		Network speed elasticity [β_{SP}]	Service elasticities		Technical change [$-\beta\tau$]
Short-run [$1/\beta_y$]	Long-run [$(1-\beta_k)/\beta_y$]		Intercity [$\exp(\beta_{INTC})-1$]	Mixed [$\exp(\beta_{MIX})-1$]	
1.93 (0.18)	1.83 (0.08)	−0.26 (0.05)	−0.05 (0.02)	−0.07 (0.03)	0.014 (0.003)

[a] Estimated asymptotic standard errors are given in parentheses

cally significant. The first value, −0.05, means that a company operating in the intercity sector would suffer lower costs than an urban firm, and this probably reflects a lesser difficulty in managing networks out of the inner city. Estimated cost elasticity for the mixed service, −0.07, indicates a lower operating cost for mixed networks not only with respect to urban firms, but also compared with the intercity ones.[47] This suggests the existence of possible cost benefits, associated with the combined provision of urban and intercity services, which could arise from the better saturation of some sharable inputs.[48]

Finally, Table 6 presents the estimated rate of Hicks neutral technological change, i.e. the rate of cost diminution from 1993 to 1999 ($-\partial \ln VC_{ft}/\partial\tau_{ft} = -\beta_\tau$).[49] 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%. 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,[50] made possible by the generous grants-in-aid government programs.

[47] The differential impact on variable costs due to the shift of the firm's production from intercity to mixed services can be calculated as $[\exp(\beta_{MIX} - \beta_{INTC}) - 1]$ and is equal to −0.02.

[48] In particular, we consider the workforce (drivers and administrative staff) and, perhaps to a lesser extent, the rolling stock.

[49] See Caves et al. (1981).

[50] Between 1993 and 1996 the fuel efficiency (kilometer run per liter of fuel) of the LPT companies included in our sample increased on average from 2.5 to 2.8. Furthermore, the total expenditure for spares and repairs decreased on average by about 154,000 litre per vehicle between 1996 and 1999. This probably contributed to the cost reduction over time highlighted above.

Cost inefficiency and effects of regulatory schemes

Using FRONTIER 4.1 we obtained estimates of the X-inefficiency defined by expression (6) for each LPT company in each period of observation. The mean overall cost inefficiency, that is, the arithmetic average of the estimated individual cost inefficiency for the sample firms over all the observations involved, is found to be 1.128. This means that, on average, the cost of production exceeds the minimum level frontier by 11.8% because of X-inefficiency. The positive coefficient for τ_{ft} in Table 4 ($\delta_\tau = 0.036$) suggests that the inefficiencies of the Italian LPT firms tended to increase throughout the seven-year period. The estimates for mean cost inefficiency by year confirm the worsening of performance over time: on average, the level of X-inefficiency increased slightly, from 12.2% in 1993 to 13.7 in 1999, with an upward swing during 1993–1996 and 1999 and a brief downward swing over the period 1997 to 1998. As mentioned in Section "Modelling inefficiency effects", the deterioration of cost efficiency during the first half of the nineties may be traced 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 1997–1998 could be linked to expectations of tighter financial constraints triggered by the promulgation of the reform Law n. 549 in 1995, whereas the new rise in X-inefficiency observed in the year 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 transit companies over time, individual predicted values vary considerably among firms in each year. This leads to an investigation into the role played by the other z-factors included in model (5) that, jointly with time, determine such a variability in the inefficiency levels.

Our primary concern, in this work, is with the differential impact of regulatory schemes on cost efficiency.[51] From Table 4, the negative sign of $\delta_R(-1.521)$, the parameter related to the subsidization mechanisms as such, without their interaction with network characteristics and time, confirms our conjecture of lower X-inefficiency levels for the units run under fixed-price schemes. Indeed, when compared over time, the results indicate a tendency of mean cost inefficiency to diminish for companies facing a transition from cost-plus to fixed-price reimbursement mechanisms. The differential impact of regulation is clearly observable in cases where the subsidization practice shifted from a cost-plus to a fixed-price scheme in 1996; evidence in such a direction is also found for the municipalities which changed the reimbursement mechanism in 1999.[52] On the other hand, firms always regulated by cost-plus contracts show a mean cost inefficiency increasing over the whole period.[53] Furthermore, the negative sign of the coefficient associated with the interaction of regulation with time ($\delta_{R\tau} = -0.014$) reveals the presence of a dynamic effect of fixed-price schemes, which indicates a more strong power of reducing inefficiency as more years elapsed under this type of regulation.[54]

Looking at individual inefficiency levels, 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.[55] This is due to the fact that these inefficiency estimates 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 4 shows that an increase in the network speed tends to lower X-inefficiency ($\delta_{SP} = -0.468$), as the transit company faces more favourable exogenous operating conditions, and this effect is stronger for the units subjected to fixed-price schemes ($\delta_{RSP} = -0.896$), presumably because of the higher cost reducing effort exerted by managers in the allocation of productive resources under this type of regulation. As explained in Section "modeling in efficiency effects", 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 extreme circumstances is no longer perceptible. Thus the greater efficiency recovery for some of the companies that moved towards fixed-priced mechanisms can be partially attributed to better network characteristics, as reflected in the higher level of average commercial speed. 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, are possibly imputable to worsened operating conditions, i.e. lower network speed.

So far we have focused on the differential impact of regulatory schemes over time, by comparing predicted X-inefficiencies 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

[51] Regarding this point, it is important to point out that before the estimation of the stochastic cost frontier model (3)–(5), we proceeded with conventional regression estimations of the cost function (3) with the dummy variable δ_R included in the specification. The results from a standard OLS model revealed that the introduction of fixed-price contracts lowers operating costs by about 4% on average compared to the traditional cost-plus regulation. Furthermore, to control for possible endogeneity in the contract variable δ_R (see Dalen & Gomez-Lobo 2003), we estimated a second model including firm-specific effects. The results showed that, even controlling for firm-specific effects, the shift to a subsidization practice based on fixed-price schemes still has a significant effect (although smaller than in the OLS model), lowering costs by 2.4%. With these findings in mind we proceeded to estimate the stochastic cost frontier model (3)–(5), in order to obtain more precise information about the impact on X-efficiency of fixed-price regulation.

[52] For the former group inefficiency decreased, on average, from 9.1% in 1995 to 7.3 in 1999, while for the latter on can observe a reduction from 14.4% in 1998 to 13.7 in 1999.

[53] For these productive units inefficiency increased, on average, from 12.7 in 1993 up to 17.1 in 1999.

[54] A similar effect have been found in Dalen and Gomez-Lobo (2003) for transit firms run under yardstick type contracts.

[55] Estimates of cost inefficiency for each firm in each year are available by the author on request.

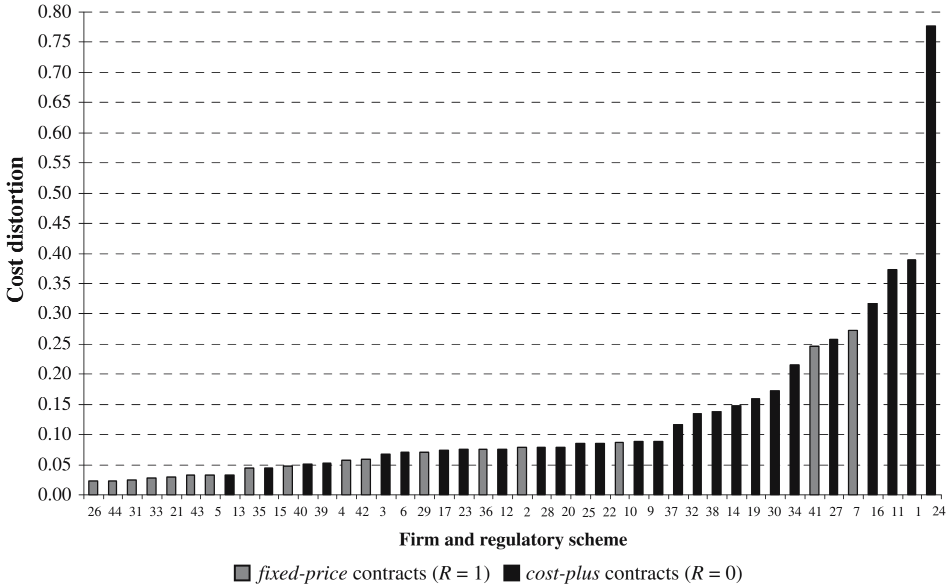

Fig. 1 Individual cost distortions over the frontier (mean 1996–1998)

the relevance of their mutual interaction, it may be convenient to fix 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 can be observed. We concentrate on the individual predicted inefficiencies pertaining to years 1996, 1997 and 1998. We have chosen this sub-period since all companies of our sample are univocally characterized by a definite regulatory mechanism during these 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 distribution of mean inefficiencies by firm and subsidization mechanism is presented in the histogram of Fig. 1. Instead of reporting $\widehat{CI}_{ft}$ we computed the percentage increase in costs due to X-inefficiency from the expression $\{\widehat{CI}_{ft}-1\}$, so the values on the vertical axis in Fig. 1 can be directly interpreted as mean cost distortions over the frontier between 1996 and 1998. They have been ranked from the best performance (Firm 26), characterized by observed operating costs that are, on average, only 2% above the frontier, to the worst performance (Firm 24), for which the cost distortion reaches 77%. An important result emerging from Fig. 1 is that 9 of the top 11 firms were subjected to fixed-price mechanisms, whereas 9 of the bottom 11 companies faced a cost-plus regulation. Once again, our findings tend to corroborate the theoretical argument by new regulatory economics about the efficacy of high powered incentive schemes in increasing efficiency.

To better understand the role played by network characteristics in the above ranking, we computed the marginal effect of the regulatory schemes and commercial speed on firms' cost distortion using the estimated parameters from the stochastic frontier and efficiency models.[56] These values, evaluated at the sample means for output, capital, input prices, and time ($\tau = 1996$) are listed in Tables 7 (*regulation impact*) and 8 (*network speed impact*), together with the percentage decrease in X-efficiency attainable by shifting from cost-plus to fixed-price regimes and by slightly improving operating conditions of the network. In order to take into account the interaction between subsidization mechanisms and network characteristics, the marginal impact of

[56] For the derivation of marginal effects see Frame and Coelli (2001) and Wang (2002).

Table 7 Impact of regulatory schemes (R) on inefficiency

Network speed[a]	Cost distortion			
	$R = 0$	Marginal effect	$R = 1$	Percentage impact
Low SP	0.152	−0.037	0.115	−24.28
Average SP	0.116	−0.047	0.069	−40.24
High SP	0.079	−0.051	0.028	−63.72

[a] Network speed levels considered in the computation of marginal effect of regulatory schemes are: low SP = 1st quartile (16 kms/h); average SP = sample mean (23 kms/h); high SP = 3rd quartile (27 kms/h)

Table 8 Impact of network speed (SP) on inefficiency

Regulatory scheme	Cost distortion			
	Average SP[a]	Marginal effect	Average SP + 1%	Percentage impact
$R = 0$	0.116	−0.015	0.101	−12.85
$R = 1$	0.069	−0.025	0.044	−36.18

[a] Average SP = sample mean level (23 kms/h)

regulatory schemes ($\partial\widehat{CI}/\partial R$) was computed for different levels of commercial speed (corresponding to sample 1st quartile, mean, and 3rd quartile), while the marginal impact of network speed ($\partial\widehat{CI}/\partial \ln SP$) was evaluated in presence of both cost-plus and fixed-price contracts.

First of all, the entries in Tables 7 and 8 confirm that both network characteristics and regulatory constraints considerably matter in determining X-efficiency of LPT firms: for a company facing an average level of commercial speed, the introduction of high powered incentive schemes allows an efficiency recovery around 40%; similarly, more favorable traffic conditions for the LPT vehicles always imply lower inefficiencies, with reductions ranging from about 13% ($R = 0$) up to 36% ($R = 1$) according to the subsidization mechanism the firm is subjected to. Secondly, a general tendency emerges for the regulation effect to become stronger as we move towards higher speed levels: for the operators characterized by a network speed higher than the sample mean, the favorable operating conditions combined with a fixed-price regulation leads to a remarkable inefficiency decrease, around 64% (from 7.9 to 2.8%); in presence of a lower speed instead, as the intrinsic technical inefficiency is likely to be rather high, the more intensive effort activity provided by managers in case of fixed-price schemes has a moderate impact on the X-efficiency, around 24%, and the global cost distortion over the frontier remains heavy (11.5%).

These results can help explain the distribution of inefficiencies in Fig. 1. It should be more clear why the top 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 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 good performances (Firm 13 and Firm 15), due to the favorable characteristics of their network, and the positioning of operators constrained by fixed-price schemes among the worst ten positions (Firm 41 and Firm 7), because of the very low levels of their commercial speed.

Conclusion and policy implications

On the whole, the results of this study indicate a significant impact of regulatory constraints on the X-efficiency of the public transit companies. First, the theoretical prediction of new regulatory economics (Laffont and Tirole 1993) that fixed-price schemes provide more incentives for efficiency is validated: given similar network characteristics, operators run under a fixed-price mechanism exhibit a lower cost distortion than operators subjected to a cost-plus regulation. Moreover, to some extent the inefficiency differentials among companies can be due to differences in the commercial speed levels. The latter contribute to determining the intrinsic inefficiency of a network and can seriously undermine the efficacy of incentive regulatory policies. In the light of the evidence found for the operators facing very low commercial speed levels, if the exogenous operating conditions are too unfavorable, then fixed-price subsidization mechanisms become less successful instruments for recovering cost efficiency.

Besides confirming the importance of incentive regulation theory for the cost analysis of public utilities, our findings also provide useful guidelines for the policy concerning local mobility. Significant reductions of X-inefficiency can be obtained by resorting to fixed-price subsidies, and the ongoing reform in Italy is correctly moving towards this direction. A proper definition of quality and cost standards is requested, so that the service contract between the regulatory authority and LPT operator gives the firm's manager the incentives to optimize the allocation of productive resources. Our results also stress the impact of network characteristics and underline the importance of local traffic regulation. 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 cost efficiency levels (higher commercial speeds move firm performances closer to the best-practice behavior). This aim could be pursued directly, 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 indirectly, through the provision of incentives for the use of public modes (e.g., higher service frequency and route density, good timetable coordination, introduction of multi-modal travelcards).[57]

In conclusion, there is a scope for public transport policy to increase the cost efficiency of LPT companies. Efforts have to be intensified in the two-fold 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 transit firms.

[57] For more details on this point see Fitzroy and Smith (1999) and Fraquelli et al. (2004).

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