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by
Anne-Kathrin Last and Heike Wetzel

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The Efficiency of German Public Theaters: A Stochastic Frontier Analysis Approach

Anne-Kathrin Last * Heike Wetzel †

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In recent years the economic performance of public non-profit sectors such as cultural services has become an interesting economic issue. This is due to the high dependence of cultural institutions on public funding on the one hand and the increasing cost-pressure on public budgets on the other hand. In order to achieve an efficiently, cost-minimizing resource allocation public authorities who decide on the distribution of public budgets need reliable performance indicators. Against this background, this paper analyzes the efficiency of German public theaters for the seasons 1991/1992 to 2005/2006. Using a stochastic frontier analysis approach, we test whether the assumption of cost-minimizing behavior is reliable in this sector. Moreover, several panel data models that differ in their ability to account for unobserved heterogeneity are applied to evaluate the impact of unobserved heterogeneity on the efficiency estimates. The results indicate that the cost-minimizing assumption cannot be maintained. Consequently, an efficiency analysis based on a cost function approach seems inappropriate in the case of German public theaters. Further, we find a considerably unobserved heterogeneity across the theaters, which causes a significant variation in the models' efficiency estimates. This implies that failing to account for unobserved heterogeneity leads to biased efficiency values. Overall, our results suggest that there is still space for improvement in the employment of resources in the sector.

Keywords: Public theaters, efficiency, stochastic frontier analysis, input distance function, cultural economics

JEL-Classification: L82, D24, Z10

*Leuphana University Lueneburg, Institute of Economics, P.O. Box 2440, 21314 Lüneburg, Germany, Phone: +49-4131-677-2301, Fax: +49-4131-677-2026, e-mail: last@uni.leuphana.de

†Leuphana University Lüneburg, Institute of Economics, P.O. Box 2440, 21314 Lüneburg, Germany, Phone: +49-4131-677-2324, Fax: +49-4131-677-2026, e-mail: wetzel@uni.leuphana.de

1 Introduction

Like in most other European countries, the amount of public funding of cultural institutions in Germany is relatively high. This is due to national cultural policy aims such as providing a cultural infrastructure and preserving cultural heritage. In particular, performing arts institutions with cost-intensive production processes and relatively low revenues rely heavily on public funding; on average, 81 percent of the income of German public theaters comes from subsidies granted by local, federal or European public authorities (Deutscher Bühnenverein (German Stage Association), 2008).

Given this high level of public funding and the increasing cost-pressure on public budgets, the economic performance or efficiency of German public theaters has become an interesting economic issue. In particular, public authorities who decide on the distribution of public budgets need reliable performance indicators in order to promote an optimal, cost-minimizing employment of resources. In this context, the application of benchmarking methods in public non-profit sectors such as higher education, health care or cultural services has become more important in recent years. Benchmarking or efficiency analysis methods compare the economic performance of an individual firm to a reference set of firms; that is, they detect the "best-practice" firms and provide clues to the management of non-profit institutions as well as for public authorities for a more efficient use of resources.

This paper gives further insights on the employment of resources for performing arts productions. In contrast to other studies in this area that have used the non-parametric data envelopment analysis method, we use the parametric stochastic frontier analysis approach to evaluate the efficiency of public theaters in Germany. For that purpose, we employ an input distance function that requires no specific behavior assumptions, such as cost-minimization or profit-maximization, and estimates the degree to which a theater could reduce its use of inputs to produce a certain level of output. Further, in order to compare our findings to those of previous studies that have utilized a cost function approach and, in particular, to test their assumption of cost-minimizing behavior, we also employ a cost function approach. To our knowledge, this is the first study on performing arts institutions that applies both an input distance function and a cost function model.

Moreover, since several studies have shown that failing to account for unobserved firm-specific heterogeneity in performance measurement can result in overestimated inefficiency values (see, e.g. Farsi et al., 2005; Greene, 2005*a*,*b*), several panel data models that differ in their ability to account for unobserved heterogeneity are applied in order to evaluate the impact of unobserved heterogeneity on the efficiency estimates. In particular, we compare the estimation results of four different stochastic frontier models for panel data: the fixed effects model of Schmidt and Sickles (1984), the random effects model of Pitt and Lee (1981), the true random effects model proposed by Greene (2005*a*,*b*), and the true random effects model with a Mundlak (1978) adjustment, as suggested by Farsi et al. (2005). The panel data set employed consists of 174 German public theaters and covers the 1991/1992 season through the 2005/2006 season.

The paper is organized as follows. Section 2 discusses theoretical aspects of the financing of German public theaters and summarizes previous research on the economic performance measurement of theaters. Specifications of the applied models are introduced in Section 3, followed by a presentation of the estimation approach in Section 4, and a description of the data in Section 5. Estimation results of the empirical analysis are presented in Section 6, and Section 7 concludes the paper.

2 Theoretical background and previous research

The current analysis investigates the efficiency of public theaters in Germany that are funded by public authorities. Agreements between the public authorities and the theaters set the amount of funds granted to the theaters for periods of time that can span several seasons, so theaters can plan the number of new productions and performances and the corresponding factor input for the coming seasons.

According to Niskanen's model of bureaucracy (1971) and its further development by Migué and Bélanger (1974), managers of organizations in the public sector try to maximize either the quantity of output or the resources granted by public authorities because these two parameters are positively correlated with the arguments in their utility functions. As Taalas proposed (1997), such arguments can be "desire for large audiences, high quality of productions, or large budgets." Thus, according to Niskanen's model, a higher output or a higher amount of available resources increases managers' utility. It is assumed that both public authorities and theater managers know the demand in terms of quantity and quality of theater productions, whereas only the managers are fully aware of the minimum costs. This information asymmetry enables theater managers to extend either the output or the requested funds beyond the market equilibrium, based on their own preferences (Migué and Bélanger, 1974).

Following the model of bureaucracy, in the case of German public theaters, where the amount of funds is fixed by the agreements between the theaters and the public authorities that fund them, theater managers can utilize the information asymmetry regarding the minimum costs in order to maximize their utility. If they claim higher than the minimum costs for a certain output level, they will have more available resources for each theater production, which enables them, for example, to hire famous actors or to indulge in complex and expensive stage designs. Given the chosen output level, this suggests that the inputs could be reduced.

However, one might argue that the assumption of information asymmetry between theater managers and public authorities cannot be sustained. The German stage association (Deutscher Bühnenverein) publishes annual statistics that provide information on the revenues and expenditures of all German public theaters. Although, these statistics provide a general overview on the production structure of the theaters they do not allow to gain comprehensive theater-specific information about the cost-minimizing input employment without a detailed and time-consuming cost structure analysis. Assuming that budget and time constraints do not allow such a detailed analysis by the public authorities in each single case, at least a partial information asymmetry remains.

Referring to the subsidy system, the promotion of a cost-minimizing employment of the granted resources is intended by the possibility to transfer unspent public funds from one season to the next and, thus, to increase the budget for the next season. However, the possibility for a higher budget in a subsequent season is a rather weak incentive for spending less than the available funds in the current season because the theater managers must fear that the subsidies will be reduced in case of recurrent savings and transfers. Altogether, considering the theoretical explanations and the actual subsidy system of public theaters in Germany, it is likely that the assumption of cost-minimizing behavior is violated.

Several empirical studies have been conducted on the cost structure of performing arts productions to assess potential economies of scale (see, e.g., Throsby, 1977; Globerman and Book, 1974).[1] These studies have all estimated a cost function, which implies that the assumption of cost-minimizing behavior holds. Moreover, most newer studies have used panel data sets to gain further information and to consider technical progress. For example, Fazioli and Filippini (1997) used a dataset of 28 Italian theaters during the period 1991-1993; to estimate a short-run cost function, they assumed that theater managements minimize the variable costs for the production of performances and found economies of scale and scope within the production process. Therefore, they suggested that established productions should be performed more often in different theaters in order to take advantage of the size effects and that, because of the economies of scope in different theater activities, theater groups should avoid specializing in particular types of performances.

Taalas (1997) was among the first to address possible inefficiencies regarding the employment of inputs in performing arts productions. He analyzed the cost structure of 37 Finnish theaters during 1985-1993 and found, as in the other studies, that the cost function estimates indicated economies of scale in production. In order to test for allocative efficiency, Taalas compared the estimated shadow prices with the market prices and found significant differences, so he rejected the assumption of allocative efficiency with respect to the use of inputs. His estimation results showed that, on average, the observed total costs exceeded the minimum costs by nearly 5 percent. Taalas also suggested that the violation of the cost-minimizing behavior may have been caused either by the restrictions implied through public subsidies or by rent-seeking by theater managers. Therefore, as Taalas noted, "a need for a detailed analysis of the extent of technical inefficiencies is accentuated. A full blown examination of both allocative and technical efficiencies in the production of cultural services, however, necessitates the application of cost or production frontiers."

So far, only a few studies have used frontier techniques to identify possible sources of inefficiencies in the production process of performing arts. Marco-Serrano (2006) measured the technical efficiency of an unbalanced panel of Spanish theaters organized in a network in the Valencia region during 1995-1999. His results, obtained by means of data envelopment analysis, showed a decrease in the efficiency scores over the analyzed

[1] For a detailed overview, see Marco-Serrano (2006).

period, over which the network expanded continuously because of the incorporation of new theaters.

Another study that applied data envelopment analysis was conducted by Tobias (2003), who analyzed the cost efficiency of German Public Theaters for the seasons 1995/96-1998/99 using the same data source as in the present article. His overall findings reported average cost inefficiencies of about 11 percent.

Overall, previous studies on the cost structure of public theaters have focused on economies of scale and scope, although more recent work has concentrated on possible inefficiencies in production as a result of the finding that the assumption of cost-minimizing behavior is likely to be violated. Thus far, data envelopment analysis has been applied to measure the inefficiencies in the context of performing arts productions, but the current study uses a stochastic frontier analysis approach for panel data that accounts for stochastic influences within the data.

3 Model specification

In order to analyze the economic performance of German public theaters, we apply an input distance function approach. Compared to a cost function approach, the input distance function approach requires no preimposed behavioral assumption, such as cost-minimization, which is likely to be violated in the case of the highly subsidized German public theater sector. Nevertheless, in order to compare our findings to previous studies that have utilized a cost function approach and in order to test their assumption of cost-minimizing behavior, we also employ a cost function approach. If our hypothesis that German public theaters do not show a cost-minimizing behavior is supported, the coefficient estimates of both functions must differ because the linkage between the cost and the distance function, based on the duality theory, would be lost.

By modeling a production technology as an input distance function, one can investigate how much the input vector can be proportionally reduced while holding the output vector fixed. Following Coelli et al. (2005), the input distance function can be defined as:

$$D_I\left(x, y\right) = \max\{\theta : (x/\theta) \in L\left(y\right)\}, \tag{1}$$

where $L(y)$ represents the set of all non-negative input vectors $x = (x_1, ..., x_K) \in \mathbb{R}_+^K$ that can produce the non-negative output vector $y = (y_1, ..., y_M) \in \mathbb{R}_+^M$; and θ measures the proportional reduction of the input vector x. The function is homogenous of degree one in inputs and satisfies the economic regularity conditions of monotonicity and concavity, that is, the function is non-decreasing and concave in inputs and non-increasing in outputs (Kumbhakar and Lovell, 2000).

From $x \in L(y)$, $D_I(x, y) \geq 1$ follows. A value equal to one identifies the respective input vector x as being fully efficient and located on the frontier of the input set. Values greater than one belong to inefficient input vectors above the frontier. This concept is closely related to Farell's (1957) measure of input-oriented technical efficiency, which can be calculated by the reciprocal of the input distance function:

$$TE(x, y) = 1/D_I\left(x, y\right) \leq 1. \tag{2}$$

Technical efficiency values equal to one identify efficient firms using an input vector located on the production frontier. Technical efficiency values between zero and one belong to inefficient firms using an input vector above the frontier.

To estimate the input distance function we adopt a translog (transcendental-logarithmic) functional form. Unlike a Cobb-Douglas form, which assumes the same production elasticities, the same scale elasticities, and a substitution elasticity equal to one for all firms, the translog does not impose such restrictions, so it is more flexible (Coelli et al., 2005).

The translog input distance function for K (k=1,...,K) inputs and M (m=1,...,M) outputs can be written as

$$\begin{aligned} ln\, D_{it}^{I} &= \alpha + \sum_{m=1}^{M} \alpha_m \, ln\, y_{mit} + \frac{1}{2} \sum_{m=1}^{M} \sum_{n=1}^{M} \alpha_{mn} \, ln\, y_{mit} \, ln\, y_{nit} + \sum_{k=1}^{K} \beta_k \, ln\, x_{kit} \\ &+ \frac{1}{2} \sum_{k=1}^{K} \sum_{l=1}^{K} \beta_{kl} \, ln\, x_{kit} \, ln\, x_{lit} + \sum_{k=1}^{K} \sum_{m=1}^{M} \gamma_{km} \, ln\, x_{kit} \, ln\, y_{mit} \\ &+ \theta_t \, t + \frac{1}{2} \theta_{tt} \, t^2 + \sum_{k=1}^{K} \lambda_{kt} \, ln\, x_{kit} \, t + \sum_{m=1}^{M} \phi_{mt} \, ln\, y_{mit} \, t + \sum_{s=1}^{S} \psi_s \, z_{sit}, \end{aligned} \tag{3}$$

where the subscripts i and t denote the firm and year, respectively; D_{it}^{I} is the input distance term; x_{kit} and y_{mit} denote the input and output quantity, respectively; $t = 1, ..., T$ is a time trend; z_{sit} $(z = 1, ..., S)$ is a vector of observable firm-characteristics expected to influence the production technology; and $\alpha, \beta, \gamma, \theta, \lambda, \phi$, and ψ are unknown parameters to be estimated.

For the theoretical conditions of symmetry and linear homogeneity in inputs to be guaranteed, several linear restrictions must hold for the input distance function. Symmetry requires the restrictions

$$\alpha_{mn} = \alpha_{nm}, \quad (m, n = 1, 2, ..., M) \quad \text{and} \quad \beta_{kl} = \beta_{lk}, \quad (k, l = 1, 2, ..., K), \tag{4}$$

and linear homogeneity in inputs is given if

$$\sum_{k=1}^{K} \beta_k = 1, \quad \sum_{l=1}^{K} \beta_{kl} = 0, \quad \sum_{k=1}^{K} \gamma_{km} = 0, \quad \text{and} \quad \sum_{k=1}^{K} \lambda_{kt} = 0. \tag{5}$$

Imposing the homogeneity restrictions by normalizing the distance term and the inputs in Equation 3 by one of the inputs (Lovell et al., 1994), and replacing the negative log of the distance term $-ln\, D_{it}^{I}$ with an error term ϵ_{it}, yields the estimable form of the translog input distance function.[2] The function can be written as

$$
\begin{aligned}
-\,ln\, x_{Kit} \;=\;& \alpha + \sum_{m=1}^{M} \alpha_m \,ln\, y_{mit} + \frac{1}{2}\sum_{m=1}^{M}\sum_{n=1}^{M} \alpha_{mn}\,ln\, y_{mit}\,ln\, y_{nit} + \sum_{k=1}^{K-1} \beta_k \,ln\, x_{kit}^{*} \\
&+ \frac{1}{2}\sum_{k=1}^{K-1}\sum_{l=1}^{K-1} \beta_{kl}\,ln\, x_{kit}^{*}\,ln\, x_{lit}^{*} + \sum_{k=1}^{K-1}\sum_{m=1}^{M} \gamma_{km}\,ln\, x_{kit}^{*}\,ln\, y_{mit} \\
&+ \theta_t\, t + \frac{1}{2}\theta_{tt}\, t^2 + \sum_{k=1}^{K} \lambda_{kt}\,ln\, x_{kit}^{*}\, t + \sum_{m=1}^{M} \phi_{mt}\,ln\, y_{mit}\, t + \sum_{s=1}^{S} \psi_s\, z_{sit} + \epsilon_{it},
\end{aligned} \tag{6}
$$

where $x_{kit}^{*} = (x_{kit}/x_{Kit})$.

If it is assumed that the firms follow a cost-minimization behavior and that they take input prices as given, economic theory implies duality between the input distance function and the cost function approach. Following Shephard (1953), the dual cost function can be defined as:

$$C\,(x, y) = \min_{x}\{w'x : x \in L\,(y)\} = \min_{x}\{w'x : D_I(x, y) \geq 1\}, \tag{7}$$

where $L(y)$, x and y are as defined above; and w' is a strictly positive input price vector, $w = (w_1, ..., w_K)' \in \mathbb{R}_{++}^{K}$. The function is non-negative and homogenous of degree one in input prices and satisfies the economic regularity conditions of monotonicity and concavity, that is, the function is non-decreasing and concave in input prices and non-decreasing in outputs (Färe and Primont, 1995).

Using a translog functional form, the cost function in Equation 7 can be written as:

$$
\begin{aligned}
ln\, C_{it} \;=\;& \alpha + \sum_{m=1}^{M} \alpha_m \,ln\, y_{mit} + \frac{1}{2}\sum_{m=1}^{M}\sum_{n=1}^{M} \alpha_{mn}\,ln\, y_{mit}\,ln\, y_{nit} + \sum_{k=1}^{K} \beta_k \,ln\, w_{kit} \\
&+ \frac{1}{2}\sum_{k=1}^{K}\sum_{l=1}^{K} \beta_{kl}\,ln\, w_{kit}\,ln\, w_{lit} + \sum_{k=1}^{K}\sum_{m=1}^{M} \gamma_{km}\,ln\, w_{kit}\,ln\, y_{mit} \\
&+ \theta_t\, t + \frac{1}{2}\theta_{tt}\, t^2 + \sum_{k=1}^{K} \lambda_{kt}\,ln\, w_{kit}\, t + \sum_{m=1}^{M} \phi_{mt}\,ln\, y_{mit}\, t + \sum_{s=1}^{S} \psi_s\, z_{sit},
\end{aligned} \tag{8}
$$

where C_{it} are the total costs; w_k denotes the k-th input price (k=1,...,K); and all other variables and parameters are as defined above.

Just as for the input distance function, the restrictions defined in Equation 4 and 5 must be valid in order to guarantee symmetry and linear homogeneity in input prices of the cost function. Imposing the homogeneity restrictions by normalizing total costs and

[2] The symmetry restrictions in Equation 4 are imposed during estimation.

the input prices in Equation 8 by one of the input prices, and adding an error term ϵ_{it}, yields the estimable form of the translog cost function. The function can be written as:

$$\begin{aligned}
-\,ln\,C^*_{it} \;=\;& \alpha + \sum_{m=1}^{M} \alpha_m\, ln\, y_{mit} + \frac{1}{2}\sum_{m=1}^{M}\sum_{n=1}^{M} \alpha_{mn}\, ln\, y_{mit}\, ln\, y_{nit} + \sum_{k=1}^{K-1} \beta_k\, ln\, w^*_{kit} \\
& + \frac{1}{2}\sum_{k=1}^{K-1}\sum_{l=1}^{K-1} \beta_{kl}\, ln\, w^*_{kit}\, ln\, w^*_{lit} + \sum_{k=1}^{K-1}\sum_{m=1}^{M} \gamma_{km}\, ln\, w^*_{kit}\, ln\, y_{mit} \\
& + \theta_t\, t + \frac{1}{2}\theta_{tt}\, t^2 + \sum_{k=1}^{K} \lambda_{kt}\, ln\, w^*_{kit}\, t + \sum_{m=1}^{M} \phi_{mt}\, ln\, y_{mit}\, t + \sum_{s=1}^{S} \psi_s\, z_{sit} + \epsilon_{it},
\end{aligned} \tag{9}$$

where $C^*_{it} = (C_{it}/w_{Kit})$ and $w^*_{kit} = (w_{kit}/w_{Kit})$.[3]

In addition to the homogeneity and symmetry restrictions, the input distance function and the cost function must satisfy the regularity conditions of monotonicity and concavity. If these conditions are violated, the estimated parameters are not consistent with economic theory and, hence, are not reliable (Sauer et al., 2006). For example, a monotonicity-violating input distance function will provide incorrectly signed elasticity estimates and implies the economically absurd result that, for fixed outputs, an increase in inputs will improve economic performance (O'Donnell and Coelli, 2005). Further, cost function estimates that violate the regularity conditions give reason to question the assumption of cost-minimizing behavior; duality theory fails and the elasticities of the cost function with respect to input prices are not equal to the corresponding input cost shares (Rungsuriyawiboon and Coelli, 2006).

For the purpose of securing theoretical consistency, monotonicity and concavity can either be imposed ex ante on the estimated function or tested after the estimation. As Kuenzle (2005) argued, the advantage of an after-estimation test is that it provides further insight into the empirical model and the industry; that is, after-estimation tests can provide hints about whether the functional form, the variables and the assumed behavior of the firms provide an appropriate image of the industry under consideration. Since we want to investigate whether a translog cost function approach with a cost-minimizing assumption is appropriate in the case of German public theaters, we follow this argument and test the regularity conditions ex post.[4]

To be monotone, the input distance function must be non-decreasing in inputs and non-increasing in outputs at each data point. For the translog case, this condition is equivalent to the restrictions:

$$\epsilon_k = \frac{\partial lnD}{\partial lnx_k} = \beta_k + \sum_{l=1}^{K} \beta_{kl}\, ln\, x_k + \sum_{m=1}^{M} \gamma_{km}\, ln\, y_m + \lambda_{kt}\, t \geq 0 \tag{10}$$

[3] Alternative model specifications for the input distance and the cost function, such as a Cobb-Douglas functional form, a translog functional form with no technical change and a translog functional form with Hicks neutral technical change, have been tested and rejected by likelihood-ratio tests.

[4] For a method to impose regularity conditions ex ante on the estimated function, see O'Donnell and Coelli (2005).

and

$$\epsilon_m = \frac{\partial lnD}{\partial lny_m} = \alpha_m + \sum_{n=1}^{M} \alpha_{mn}\, ln\, y_m + \sum_{k=1}^{K} \gamma_{km}\, ln\, x_k + \phi_{mt}\, t \leq 0; \tag{11}$$

that is, the elasticity of D with respect to x_k is non-negative, and the elasticity of D with respect to y_m is non-positive for all inputs and outputs.

Similarly, for the cost function, monotonicity holds if the function is non-decreasing in input prices and non-decreasing in outputs at each data point. For the translog case, this condition is equivalent to the restrictions:

$$\epsilon_k = \frac{\partial lnC}{\partial lnw_k} = \beta_k + \sum_{l=1}^{K} \beta_{kl}\, ln\, w_k + \sum_{m=1}^{M} \gamma_{km}\, ln\, y_m + \lambda_{kt}\, t \geq 0 \tag{12}$$

and

$$\epsilon_m = \frac{\partial lnC}{\partial lny_m} = \alpha_m + \sum_{n=1}^{M} \alpha_{mn}\, ln\, y_m + \sum_{k=1}^{K} \gamma_{km}\, ln\, x_k + \phi_{mt}\, t \geq 0; \tag{13}$$

that is, both the elasticity of C with respect to x_k and the elasticity of C with respect to y_m are non-negative for all input prices and outputs.

Satisfying the concavity condition requires that the input distance function be concave in inputs and the cost function be concave in input prices. For the input distance function, this requires that the Hessian matrix of the second-order derivatives of D with respect to x_k be negative semi-definite; that is, all leading principle minors of the matrix must alternate in sign, beginning with negative. Similarly, the cost function is concave in input prices if negative semi-definiteness holds for the Hessian matrix of the second-order derivatives of C with respect to w_k.

Following Diewert and Wales (1987), it can be shown that the Hessian matrix of the translog input distance function with respect to the inputs – or the translog cost function with respect to the input prices – is negative semi-definite if, and only if, the matrix $\hat{H}$ is negative semi-definite. $\hat{H}$ is defined as:

$$\hat{H} = \begin{bmatrix} \beta_{11} & \cdots & \beta_{1l} \\ \vdots & \ddots & \vdots \\ \beta_{k1} & \cdots & \beta_{kk} \end{bmatrix} - \begin{bmatrix} e_1 & \cdots & 0 \\ \vdots & \ddots & \vdots \\ 0 & \cdots & e_k \end{bmatrix} + \begin{bmatrix} e_1 e_1 & \cdots & e_1 e_l \\ \vdots & \ddots & \vdots \\ e_n e_1 & \cdots & e_k e_k \end{bmatrix} \tag{14}$$

where, for the input distance function case, β_{kl} are the second-order coefficients of the inputs and e_k are the input elasticities, and, for the cost function case, β_{kl} are the second-order coefficients of the input prices and e_k are the input price elasticities.

4 Estimation approach

In order to estimate the technical and cost efficiency of German public theaters and to investigate the influence of unobserved heterogeneity on the estimation results, we follow Farsi et al. (2005) and apply four different stochastic frontier models for panel data. Following Greene (2005a), a general stochastic frontier model for panel data can be expressed as:

$$d_{it} = \alpha + \beta' r_{it} + v_{it} + S u_i, \tag{15}$$

where d_{it} and r_{it} stand either for the dependent and independent variables in the distance function model, as specified in Equation 6, or for the dependent and independent variables in the cost function model, as specified in Equation 8. Furthermore, v_{it} and u_i are the two parts of a composed random error, $\epsilon_{it} = v_{it} + S u_i$, where v_{it} represents a random error term that captures noise as well as any firm and time-specific unobserved heterogeneity; and u_i is a non-negative time-invariant firm-specific inefficiency term. S equals -1 in the case of a distance function and 1 in the case of a cost function. Finally, α is a constant and β' is a vector of coefficients to be estimated.

The four models differ in their assumptions about the distribution of the two error components, as well as in their estimation approach for the inefficiency indicators and, therefore, in their efficiency scores. Given Farrell's measure of technical efficiency and given the definition of cost efficiency as the ratio of optimal costs to observed costs, the efficiency scores range between zero and one, where a value equal to one indicates efficiency. A summary of the econometric specifications of the models is given in Table 1.

Table 1: Econometric specifications[a]

	Model I Fixed effects	Model II Random Effects	Model III True random effects	Model IV True random effects with Mundlak formulation
Firm-specific component	fixed	$u_i \sim iidN^+(0, \sigma_u^2)$	$\alpha_i \sim iidN(0, \sigma_\alpha^2)$	$\alpha_i = \gamma \bar{x}_i + \delta_i$ $\bar{x}_i = \frac{1}{T_i}\sum_{t=1}^{T_i} x_{it}$ $\delta_i \sim N(0, \sigma_\delta^2)$
Random error ϵ_{it}	$\epsilon_{it} = v_{it}$ $v_{it} \sim iid(0, \sigma_v^2)$	$\epsilon_{it} = v_{it} + S u_i$ $v_{it} \sim iidN(0, \sigma_v^2)$ $u_i \sim iidN^+(0, \sigma_u^2)$	$\epsilon_{it} = v_{it} + S u_{it}$ $v_{it} \sim iidN(0, \sigma_v^2)$ $u_{it} \sim iidN^+(0, \sigma_u^2)$	$\epsilon_{it} = v_{it} + S u_{it}$ $v_{it} \sim iidN(0, \sigma_v^2)$ $u_{it} \sim iidN^+(0, \sigma_u^2)$
Inefficiency	$max\{\hat{\alpha}_i^d\} - \hat{\alpha}_i^d$ $\hat{\alpha}_i^c - min\{\hat{\alpha}_i^c\}$	$E[u_i \vert \bar{\epsilon}_i]$	$E[u_{it} \vert \hat{w}_{it}]$ $w_{it} = \alpha_i + \epsilon_{it}$	$E[u_{it} \vert \hat{w}_{it}]$ $w_{it} = \delta_i + \epsilon_{it}$
Relative efficiency	$e^{-(max\{\hat{\alpha}_i^d\} - \hat{\alpha}_i^d)}$ $e^{-(\hat{\alpha}_i^c - min\{\hat{\alpha}_i^c\})}$	$E[e^{-u_i} \vert \bar{\epsilon}_i]$	$E[e^{-u_{it}} \vert \hat{w}_{it}]$ $w_{it} = \alpha_i + \epsilon_{it}$	$E[e^{-u_{it}} \vert \hat{w}_{it}]$ $w_{it} = \delta_i + \epsilon_{it}$

[a]The superscripts d and c stand for the distance function and the cost function, respectively.

Model I is a fixed effects model as proposed by Schmidt and Sickles (1984). Apart from the assumption that v_{it} is independent and identically distributed (iid) with zero mean and constant variance and is uncorrelated with the explanatory variables no further distributional assumptions are required. In particular, the firm-specific inefficiency term u_i and, hence, the firm-specific fixed effects $\alpha_i = \alpha + S u_i$ are allowed to correlate with the explanatory variables or with v_{it}. In case of an input distance function, the firm's inefficiency is estimated by the deviation from the firm-specific intercept $\hat{\alpha}_i^d$ to the maximum intercept in the sample, that is $\hat{u}_i^d = max\{\hat{\alpha}_i^d\} - \hat{\alpha}_i^d$, and in the case of a cost function by the deviation from the firm-specific intercept $\hat{\alpha}_i^c$ to the minimum intercept in the sample, that is $\hat{u}_i^c = \hat{\alpha}_i^c - min\{\hat{\alpha}_i^c\}$.

Model II is a random effects model as proposed by Pitt and Lee (1981). In this model the firm-specific inefficiency term u_i is an iid half-normally distributed random effect which is independently distributed from the iid normally distributed random error term v_{it}. Further, both error components are assumed to be uncorrelated with the explanatory variables and each other. The model estimates are obtained by maximum likelihood estimation, and, as proposed by Jondrow et al. (1982), the firm's inefficiency is estimated by the conditional mean of the inefficiency term $\hat{u}_i = E[u_i|\bar{\epsilon}_i]$, where $\bar{\epsilon}_i = \frac{1}{T_i}\sum_{t=1}^{T_i}\hat{\epsilon}_{it}$ (Farsi et al., 2005).

Both models have assets and drawbacks. Since the fixed effects model identifies at least one firm as 100% efficient, the inefficiency of the other firms can only be measured relative to the best-practice firm(s) in the sample. That is, the inefficiency estimates are sensitive to sample selection and outliers (Kuenzle, 2005; Farsi and Filipini, 2004). Further, since the fixed effects estimator does not allow to include any time-invariant variables in the estimation, it cannot account for any observed time-invariant heterogeneity. Both problems can be solved by the random effects model. However, the main advantage of the fixed effects model is that the estimated coefficients are not affected by any correlation between the firm-specific effects and the explanatory variables. In contrast, the estimates of the random effects model will be biased in this case.

Moreover, the fixed and random effects models have two additional shortcomings in common. First, both models assume constant inefficiency over time which is rather unrealistic in relatively long panels. Second, any time-invariant firm-specific unobserved heterogeneity is included in the firm-specific component and thus in the inefficiency estimates. In other words, if any time-invariant firm-specific unobserved heterogeneity is existent, the conventional fixed and random effects models tend to overestimate the inefficiency (Greene, 2005*a*; Farsi et al., 2005).

Model III is the true random effects model proposed by Greene (2005*a*,*b*). This model accounts for the shortcomings of the conventional panel data models by adding a firm-specific random term α_i. The model can be expressed as

$$d_{it} = \alpha + \alpha_i + \beta' r_{it} + v_{it} + S u_{it}, \tag{16}$$

where α_i represents time-invariant firm-specific unobserved heterogeneity and u_{it} represents time-varying firm-specific inefficiency. All other notation is as defined before. The error components v_{it} and u_{it} are distributed as in the random effects model and the unobserved heterogeneity component α_i is iid normally distributed. Further, v_{it}, u_{it} and α_i are assumed to be uncorrelated with the explanatory variables and each other. As in the random effects model, the firm's inefficiency is estimated by the conditional mean of the inefficiency term $\hat{u}_{it} = E[u_i|\hat{w}_{it}]$, where $w_{it} = \alpha_i + \epsilon_{it}$.

This specification separates time-invariant unobserved heterogeneity from time-varying inefficiency and therefore relaxes the main limitations of the conventional panel data models. However, since all time-invariant effects are treated as unobserved heterogeneity and are captured by the firm-specific constant any persistent inefficiency is not included in the inefficiency term. Consequently, as the conventional fixed and random effects models tend to overestimate the inefficiency, the true random effects model tends to underestimate it (Farsi et al., 2006).

Model IV is an extension to Model III. As suggested by Farsi et al. (2005), it uses Mundlak's (1978) formulation to overcome the possible bias problem in the true random effects model that results if any correlation between the unobserved heterogeneity component α_i and the explanatory variables r_{it} is existent. Mundlak's formulation accounts for these correlations with an auxiliary regression that can be written as:

$$\alpha_i = \gamma' \bar{r}_i + \delta_i, \quad \bar{r}_i = \frac{1}{T_i} \sum_{t=1}^{T_i} r_{it}, \quad \delta_i \sim N(0, \sigma_\delta^2), \tag{17}$$

where $\bar{r}_i$ represents a vector of the group means of the explanatory variables r_{it}; γ' is the corresponding vector of coefficients to be estimated; and δ_i is an normally distributed random term that is not correlated with the explanatory variables. Incorporated in Equation 16 the auxiliary coefficients γ_x capture any linear correlation between α_i and $\bar{r}_i$ and, thus, minimize the possible bias of the main model's coefficients (Farsi et al., 2005).

5 Data

The utilized data set is an unbalanced panel of 174 German public theaters observed for the seasons 1991/1992 to 2005/2006. The data were taken from the theater reports published annually by the Deutscher Bühnenverein (German Stage Association) (1993-2007).

First, to identify and eliminate any outliers we apply the method suggested by Hadi (1992, 1994), which identifies multiple outliers in multivariate data. Moreover, all theaters with less than four observations are excluded from the estimation. For the input distance function model this procedure leaves a total of 1433 observations from 126 theaters and for the cost function model a total of 1413 observations from 124 theaters, respectively. For both models, we use a supply-based output measure as proposed by Tobias (2003) and include three input variables as measures for labor and capital input.

Since the theaters run stages with auditoriums of different sizes, including only the number of performances as an output measure would bias the use of inputs regarding the quantity of output. Therefore, in order to account for the differences in size, we measure the output using the variable *number of supplied tickets*, calculated as the number of performances per season multiplied by the number of seats.[5] For the distance function model, the total salary expenses and the *operating expenses* per season are used as

[5] Most theaters run several stages so, the number of supplied tickets is calculated for every stage and then summed.

monetary measures for the quantities of labor and capital.[6] The salary expenses are divided into *salary expenses for artistic staff* and *salary expenses for administrative and technical staff*. This division allows for a more detailed identification of possible sources of inefficiency. The operating expenses include, among other things, administration costs, leasing and fire service expenditures.

In addition, to account for observed heterogeneity three firm characteristic variables are taken into account. First, as described in Section 2, the theaters are aware of the amount of subsidies granted by the public authorities when they plan productions for upcoming seasons, so we include a variable reflecting the *amount of subsidies* in the model in order to test for the impact of public funding on efficiency. We assume that this variable has a negative impact on efficiency. Further, the production technology of the theaters differs significantly by the number of stages that belong to one theater. Given the same amount of output, it can be assumed that a theater with more than one stage has higher input requirements than a theater with only one stage. Therefore, we expect the second firm characteristic variable included in the model, the *number of stages*, to have a negative impact on efficiency. Finally, the third firm characteristic incorporated in the model is the number of different productions per season. Besides their public mission regarding the maintenance of cultural diversity, theaters have incentives to offer a range of different plays in order to attract large audiences. However, producing plays is cost-intensive and is expected to have a negative impact on efficiency, while re-runs of an established production are much less expensive for the theater. Moreover, the necessary rearrangements of stage designs that result from changing productions, irrespective of whether the productions are new or not, result in higher costs. Therefore the variable, *number of productions* per season, is included in order to control for the impact on input requirements.

For the cost function model, in addition to the *total costs*, given as the sum of total salary expenses and operating expenses, the input variables must be included in terms of prices. Therefore, to calculate the prices for labor, the two salary expense variables are divided by the appropriate number of employees in order to get the *price for artistic staff* and the *price for administrative and technical staff*. Accordingly, following Taalas (1997), the *price for capital* is given by the operating expenses divided by the number of seats, which is used as a proxy for the capital stock. Finally, the same output and firm characteristic variables as in the input distance function model are included.

Tables 2 and 3 report the summary statistics of the output variable, the input variables and the firm characteristic variables of the input distance function model and the cost function model. The descriptive statistics show significant variance regarding all variables. For example, the largest theater in terms of output supplies 97 times more tickets than the smallest theater in the data set.[7] This variance results from the different auditorium sizes and number of stages run by each theater.

[6] All monetary measures are adjusted for inflation using the consumer price index for Germany (Statistisches Bundesamt (Federal Statistical Office), 2009). Values are stated in year-2005 Euros.

[7] The largest theater in terms of tickets supplied is *Niedersaechsisches Staatstheater Hannover*, which includes the state opera house and the *Schauspielhaus*, resulting overall in about 2360 seats. The smallest theater is the *Schlosstheater Moers*, which has about 300 seats.

Table 2: Input distance function – summary statistics

Variable description	Variable	Mean	Median	Std. Dev.	Min	Max
Number of supplied tickets (10^3)	Y	171	160	108	6	596
Salary expenses for artistic staff (10^3 EUR)	X_{Lart}	6432	5899	5082	163	27800
Salary expenses for administrative and technical staff (10^3 EUR)	X_{Lad}	5257	4229	4139	23	25300
Operating expenses (10^3 EUR)	X_C	2559	2074	1928	42	11400
Amount of subsidies (10^3 EUR)	SUB	13000	11500	9540	218	54700
Number of stages	ST	4	4	2	1	13
Number of productions	$PROD$	29	26	14	2	79
Number of observations			1433			

Source: Deutscher Bühnenverein (German Stage Association) (1993-2007)

Table 3: Cost function – summary statistics

Variable description	Variable	Mean	Median	Std. Dev.	Min	Max
Number of supplied tickets (10^3)	Y	175	163	110	6	596
Total annual costs (10^3 EUR)	TC	14600	12400	10800	326	61100
Price for artistic staff (10^3 EUR/employee)	P_{Lart}	54	51	17	13	138
Price for administrative and technical staff (10^3 EUR/employee)	P_{Lad}	37	37	9	8	76
Price for capital (EUR/seat)	P_C	1861	1534	1215	93	9851
Amount of subsidies (10^3 EUR)	SUB	13300	12000	9814	218	56000
Number of stages	ST	4	4	2	1	13
Number of productions	$PROD$	30	27	14	2	90
Number of observations			1413			

Source: Deutscher Bühnenverein (German Stage Association) (1993-2007)

Table 4 presents the fraction of within variation of the overall variation for the variables included in the two models. The figures indicate that the variables show a significant fraction of within variation. Altogether, the descriptive statistics indicate that the used variables show a reasonable between and within variation, which finding supports the use of panel data models and the use of fixed effects models in particular.

Table 4: Fraction of within variation[a]

Input distance function model		Cost function model	
Variable	Fraction of within variation	Variable	Fraction of within variation
$-ln\,(X_C)$	0.20	$ln\,(TC/P_C)$	0.32
$ln\,Y$	0.17	$ln\,Y$	0.17
$ln\,(X_{Lart}/X_C)$	0.39	$ln\,(P_{Lart}/P_C)$	0.55
$ln\,(X_{Lad}/X_C)$	0.47	$ln\,(P_{Lad}/P_C)$	0.51
$ln\,SUB$	0.11	$ln\,SUB$	0.10
$ln\,ST$	0.51	$ln\,ST$	0.53
$ln\,PROD$	0.43	$ln\,PROD$	0.45

[a] Within variation represents the standard deviation of theater observations from the theater's average $(X_{it} - \bar{X}_i)$. The fraction of within variation is defined as the ratio of within to overall standard deviation (Farsi et al., 2005).

6 Results

The parameter estimates of the input distance and the cost function for all models are presented in Table 5. As each variable is in natural logarithm and is normalized by its sample median, the first-order coefficients can be interpreted as distance and cost elasticities at the sample median firm, respectively.

Considering the different model specifications a Hausman test conducted on both functions indicates that the firm-specific effects are correlated with the explanatory variables. This suggests that the estimated coefficients of Models II and III, which do not account for this correlation, are biased. In contrast, since our data set shows a reasonable within variation and covers a sufficient long time period of 15 years, the results of the fixed effects model can be considered as unbiased parameter estimates.[8]

First, focusing on the distance function estimates, the results show that all first-order coefficients are statistically significant at the 1 percent level and have the expected signs across all models. In other words, at the sample median firm, the estimated input distance function is decreasing in outputs and increasing in inputs. The magnitude of the coefficients differs slightly across the models, indicating slightly biased estimation results for Models II and III. The smallest difference among the models' coefficients is observed between Models I and IV. In conjunction with the statistical significance of 17 out of the 20 Mundlak terms in Model IV this suggests that the applied Mundlak formulation is able to account for correlations between the firm-specific effects and the explanatory variables, and, thus, to reduce the resulting bias.[9]

8 In short panels the so called 'incidental parameter' problem arises, yielding inconsistent parameter estimates.

9 The Mundlak terms of Model IV are not reported to conserve space. For both functions 17 out of the 20 Mundlak coefficients are statistically different from zero at the 5 percent level.

Table 5: Parameter estimates[a,b,c]

		Input distance function						Cost function			
Variable	Parameter	Model I	Model II	Model III	Model IV	Variable	Parameter	Model I	Model II	Model III	Model IV
Y	α_1	−0.121(−8.6)	−0.167(−13.6)	−0.166(−34.6)	−0.122(−10.5)	Y	α_1	0.211(10.6)	0.277(18.1)	0.245(37.9)	0.210(13.0)
Y^2	α_{11}	−0.032(−2.5)	−0.006(−0.7)	−0.023(−6.4)	−0.034(−3.0)	Y^2	α_{11}	−0.002(−0.1)	0.002(0.1)	0.001(0.2)	0.006(0.4)
X_{Lart}	β_1	0.490(43.0)	0.493(44.4)	0.495(117.4)	0.494(60.6)	P_{Lart}	β_1	0.399(30.7)	0.400(42.0)	0.399(60.4)	0.399(49.2)
X_{Lad}	β_2	0.372(29.5)	0.353(28.6)	0.358(58.8)	0.366(37.3)	P_{Lad}	β_2	0.451(32.2)	0.451(32.2)	0.449(68.4)	0.447(43.5)
X_C	β_3	0.138	0.154	0.147	0.140	P_C	β_3	0.150	0.149	0.152	0.154
X_{Lart}^2	β_{11}	0.165(11.3)	0.201(14.7)	0.185(24.6)	0.170(17.5)	P_{Lart}^2	β_{11}	0.242(8.0)	0.230(12.3)	0.192(13.3)	0.187(12.1)
X_{Lad}^2	β_{22}	0.118(10.5)	0.116(9.1)	0.116(15.5)	0.117(8.2)	P_{Lad}^2	β_{22}	0.120(3.7)	0.126(4.6)	0.111(5.9)	0.103(4.9)
X_C^2	β_{33}	0.037	0.069	0.047	0.035	P_C^2	β_{33}	0.074	0.076	0.085	0.086
$X_{Lart}X_{Lad}$	β_{12}	−0.123(−12.5)	−0.124(−13.9)	−0.127(−17.9)	−0.126(−13.9)	$P_{Lart}P_{Lad}$	β_{12}	−0.144(−5.1)	−0.140(−6.1)	−0.109(−6.7)	−0.102(−5.8)
$X_{Lart}X_C$	β_{13}	−0.042	−0.077	−0.058	−0.044	$P_{Lart}P_C$	β_{13}	−0.098	−0.090	−0.083	−0.085
$X_{Lad}X_C$	β_{23}	0.005	0.008	0.011	0.009	$P_{Lad}P_C$	β_{23}	0.024	0.014	−0.002	−0.001
$X_{Lart}Y$	γ_{11}	−0.027(−3.2)	−0.001(−0.1)	−0.012(−3.0)	−0.026(−2.9)	$P_{Lart}Y$	γ_{11}	0.022(1.8)	0.016(2.2)	0.017(3.0)	0.012(1.9)
$X_{Lad}Y$	γ_{21}	0.049(4.9)	0.036(3.4)	0.040(6.5)	0.047(5.0)	$P_{Lad}Y$	γ_{21}	0.017(1.3)	0.013(1.2)	0.020(2.7)	0.022(2.1)
X_CY	γ_{31}	−0.022	−0.035	−0.028	−0.021	P_CY	γ_{31}	−0.039	−0.013	−0.037	−0.034
T	θ_t	−0.004(−6.8)	−0.005(−10.1)	−0.004(−12.4)	−0.004(−9.6)	T	θ_t	−0.005(−7.3)	−0.004(−8.9)	−0.005(−12.0)	−0.005(−11.6)
T^2	θ_{tt}	0.000(−1.2)	−0.001(−2.6)	0.000(−2.5)	0.000(−1.2)	T^2	θ_{tt}	0.003(7.5)	0.003(9.5)	0.002(7.5)	0.002(5.7)
$X_{Lart}T$	λ_{1t}	−0.004(−3.9)	−0.003(−3.5)	−0.004(−5.5)	−0.004(−5.8)	$P_{Lart}T$	λ_{1t}	0.005(2.4)	0.004(2.7)	0.003(2.7)	0.003(2.3)
$X_{Lad}T$	λ_{2t}	0.008(6.8)	0.006(5.4)	0.007(8.5)	0.008(8.4)	$P_{Lad}T$	λ_{2t}	0.000(0.1)	0.001(0.4)	0.001(0.5)	0.001(0.8)
X_CT	λ_{3t}	−0.004	−0.003	−0.003	−0.004	P_CT	λ_{3t}	−0.005	−0.005	−0.004	−0.004
YT	ϕ_{1t}	0.000(−0.4)	0.001(1.7)	0.000(1.4)	0.000(−0.3)	YT	ϕ_{1t}	0.003(4.6)	0.003(6.2)	0.002(6.6)	0.003(6.05)
SUB	ψ_1	−0.622(−40.3)	−0.783(−123.6)	−0.731(−192.2)	−0.617(−68.1)	SUB	ψ_1	0.364(15.3)	0.523(36.7)	0.408(78.2)	0.313(23.3)
ST	ψ_2	−0.006(−1.0)	0.002(0.2)	−0.002(−0.5)	−0.009(−1.4)	ST	ψ_2	0.106(9.3)	0.107(11.8)	0.111(19.2)	0.119(13.4)
$PROD$	ψ_3	−0.026(−3.6)	−0.016(−1.9)	−0.017(−4.0)	−0.025(−3.7)	$PROD$	ψ_3	0.026(2.4)	0.022(2.3)	0.028(4.9)	0.033(3.4)
Constant	α	–	0.317(21.4)	0.053(15.0)	−0.036(−9.1)	Constant	α	–	−0.464(−22.5)	−0.084(−19.5)	0.000(−0.1)
Sigma	$\sqrt{\sigma_u^2+\sigma_v^2}$	–	0.364(9.5)	0.840(11.2)	0.963(8.7)	Sigma	$\sqrt{\sigma_u^2+\sigma_v^2}$	–	0.516(13.6)	1.810(12.3)	2.217(11.5)
Lambda	σ_u/σ_v	–	6.090(3.9)	0.067(51.3)	0.067(37.3)	Lambda	σ_u/σ_v	–	6.164(5.5)	0.114(43.2)	0.119(44.2)

[a] All variables are in natural logarithm and are normalized by their sample median. [b] T-ratios are reported in parenthesis. [c] All model estimates are obtained by using Limdep 9.0.

Given the unbiased estimates of Model I, the estimated input elasticities for salary expenses for artistic staff (β_1) and for salary expenses for administrative and technical staff (β_2) are found to be equal to 0.490 and 0.372, respectively. The input elasticity for operating expenses (β_3) is calculated via the homogeneity restriction presented in Equation 5 and equals 0.138. Interpreted as shadow shares, the input elasticities indicate that expenses for artistic staff account for 49.0 percent, expenses for administrative and technical staff account for 37.2 percent, and operating expenses account for 13.8 percent of total costs at the sample median firm. These values are similar to the observed cost shares at the sample median firm that account for 48.3, 34.7, and 17.0 percent, respectively.

The first-order coefficient of time (θ_t) is -0.004. Independent of the negative sign that implies regressive technical change, the fairly low magnitude suggests almost no technical change for the sample median firm in the mid year of the sample. This result can be explained by the very limited possibilities of the performing arts to benefit from labor or capital-saving technological improvements compared to other sectors (Fazioli and Filippini, 1997).

Referring to the firms' characteristics, two out of the three coefficients are significantly different from zero at the 1 percent level. First, the statistically significant and negative coefficient of subsidies (ψ_1) implies that a 1 percent increase in the amount of subsidies will increase the input requirements by 0.62 percent at the sample median firm. This result is consistent with previous research (Bishop and Brand, 2003) and confirms our expectation that public funding has a negative impact on efficiency. Furthermore, it corroborates the hypothesis that the assumption of cost-minimizing behavior is violated in the highly subsidized German public theater sector. Second, the statistically significant and negative coefficient of the number of productions (ψ_3) suggests an increase of input requirements of 2.6 percent for an additional production at the sample median firm. This result shows that the aim of cultural diversity comes at some costs.

As for the distance function estimates, the first-order coefficients of the cost function estimates are all statistically significant at the 1 percent level and have the expected signs across all models. In other words, at the sample median firm, the estimated cost function is increasing in outputs and in input prices. Again, the magnitude of the coefficients differs slightly across the models indicating slightly biased estimation results of Models II and III. However, the unbiased estimates of Model I are rather different than the estimates of the corresponding input distance function model. In particular, the cost elasticities with respect to input prices that can be interpreted as shadow shares differ significantly from the input distance function estimates. Summarized, the cost function estimates indicate that expenses for artistic staff (β_1) account for 39.9 percent, expenses for administrative and technical staff (β_2) account for 45.1 percent, and operating expenses (β_3) account for 15.0 percent of total costs at the sample median firm. Except for the operating expenses estimate, these values are quite different from the observed cost shares as well as from the distance function shadow shares.

As noted by Rungsuriyawiboon and Coelli (2006), one possible–and in our case the most likely–reason for this difference is the violation of the cost-minimization assumption.[10] This conclusion is further supported by the posteriori test results on the violations of the regularity conditions. While the monotonicity condition is violated in less than 1 percent of the observations in both the distance and the cost function, the violation rate of the curvature condition is rather high in the cost function compared to the distance function. While the curvature condition across all estimated models is violated at the maximum of about 20 percent of the observations in the distance function, the violation rate in the cost function differs from as high as 94 percent in Model I to at least 25 percent in Model IV.[11] To sum up, these results suggest that the cost-minimizing assumption is violated; thus, the cost function estimates are rather unreliable. Consequently, an efficiency analysis based on a translog cost function approach seems to be inappropriate in the case of German public theaters.

The summary statistics of the estimated technical efficiency scores are reported in Table 6.[12] As expected, the results differ considerably across the four models. In particular, the efficiency estimates of the conventional fixed and random effects models (Models I and II) are rather low compared to the efficiency estimates of the true random effects models (Models III and IV).

Table 6: Summary statistics of technical efficiency scores

	Model I	Model II	Model III	Model IV
Minimum	0.213	0.331	0.826	0.822
Maximum	1.000	0.994	0.993	0.992
Mean	0.367	0.724	0.967	0.964
Median	0.336	0.715	0.968	0.966
5 percentile	0.239	0.617	0.951	0.944

For Model I, the mean efficiency value of 0.367 implies that on average, the same output quantity could have been produced despite reducing the input usage by more than 63 percent. Model II shows a much higher mean efficiency value of 0.724, indicating a possible input reduction of about 28 percent on average. Further, the minimum efficiency values of Model I and Model II, 0.213 and 0.331, suggest an input saving potential of about 67 to 79 percent, respectively, for the most inefficient observations. These results, particularly the mean efficiency value of Model I, seem rather unrealistic and can be explained by the fact that both models assume constant inefficiency over time and do not separate firm-specific unobserved heterogeneity from inefficiency. That is, any unobserved heterogeneity that influences the firm-specific production structure is included in the efficiency scores. Moreover, since Model I identifies at least one obser-

[10] Other possible reasons noted by Rungsuriyawiboon and Coelli (2006) are measurement errors in either the input quantities or prices or an endogenous regressor problem in the distance function model. See Rungsuriyawiboon and Coelli (2006) for a further discussion.

[11] The violation rate of the curvature condition in the distance (cost) function is 20 (94) percent in Model I, 18 (67) percent in Model II, 19 (26) percent in Model III, and 20 (25) percent in Model IV.

[12] Since the cost function estimates are considered less reliable the estimated cost efficiency scores are not reported to conserve space. The results are available on request from the authors.

vation as 100 percent efficient, its efficiency estimates are sensitive to sample selection and outliers that could result in highly downward biased efficiency values.

Turning to Models III and IV that account for the shortcomings of Models I and II by assuming time-varying inefficiency and distinguishing unobserved heterogeneity from inefficiency, the mean efficiency values of 0.967 and 0.964, respectively, indicate a possible input reduction of about 3 percent on average. Combined with the minimum efficiency values of 0.826 and 0.822, respectively, that suggest an input saving potential of about 17 percent for the most inefficient observations, and the 5th percentile values of 0.951 and 0.944, respectively, which suggest that in only 5 percent of all observations the input saving potential is higher than approximately 5 percent, these results seem to be more reasonable. Further, the nearly identical efficiency results of Models III and IV indicate that the efficiency estimates are not influenced considerably by a bias resulting from any correlation between the firm-specific effects and the explanatory variables.

7 Conclusions

In this study, we analyzed the efficiency of German public theaters for the seasons 1991/1992 to 2005/2006. Based on a stochastic frontier analysis approach, we tested whether the assumption of cost-minimizing behavior is reliable and thus, whether an efficiency analysis via a cost function approach is appropriate in this sector. In addition, several panel data models that differ in their ability to account for unobserved heterogeneity were applied to evaluate the impact of unobserved heterogeneity on the efficiency estimates.

With regard to the cost-minimizing behavior, both the differences between the estimated cost function's shadow shares and the observed cost shares as well as the high violation rate of the regularity conditions in the cost function approach suggest that the assumption of cost-minimizing behavior cannot be maintained. This conclusion is supported by the theoretically consistent result that higher subsidies lead ceteris paribus to relatively higher input requirements. Consequently, an efficiency analysis based on a cost function approach that presumes a cost-minimizing behavior may provide biased efficiency estimates and, hence, seems inappropriate in the case of German public theaters.

Referring to the impact of unobserved heterogeneity on the efficiency estimates, our results are consistent with previous research (see, e.g., Farsi and Filipini, 2004; Farsi et al., 2005). We observe considerably differences between the conventional fixed effects and random effects models that do not separate firm-specific unobserved heterogeneity from inefficiency and the two true random effects models that do. Hence, our results imply that the German public theater sector is characterized by considerably unobserved heterogeneity across the theaters, which influences the theater-specific production structure. Given the rather unrealistic efficiency estimates of the conventional models, it can be assumed that these models, particularly the fixed effects model, underestimate the efficiency. Alternatively, the true random effects models treat all time-invariant effects as unobserved heterogeneity, including any persistent inefficiency; hence, they may overestimate the efficiency to a certain degree. Nevertheless, the results of the true random effects models seem to be more realistic. Both models indicate an input saving potential of about 3 percent on average and a maximum input saving potential of about 5 percent for 95 percent of all observations.

Overall, our results suggest that there is still space for improvement in the employment of resources in the German public theater sector. In particular, given the doubts about the existence of cost-minimizing behavior in the presence of the actual subsidy system, public authorities should carefully reassess the system with a particular focus on the implementation of cost-minimizing incentives.

In this context, it should be noted that our efficiency results are solely based on a technical view of performing arts production. Due to the lack of data we were not able to include any quality aspects into our analysis and, therefore, the results provide no information on the relationship between efficiency and quality. Consequently, the judgement whether the measured technical inefficiency of an individual theater is due to poor employment of the resources or to higher input requirements as a result of a higher quality level remains with the public authorities. While budget cuts are appropriate in case of a theater that reveals both a high technical inefficiency level and a poor quality level, a certain degree of technical inefficiency might be acceptable in case of a theater that provides a high quality level and, hence, should not result in budget cuts.

To sum up, our analysis provides information on one side of the story – technical efficiency – which connected with the other side of the story – quality – can help public authorities to identify the best-performing theaters and to derive the best-practice production strategies. As suggested by Farsi et al. (2005), the results of such an analysis can be used to predict an interval of necessary costs for each theater and, thus, can reduce the information asymmetry regarding minimum costs in the context of negotiations between theater managers and public authorities.

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