The real U.S. dollar exchange rate is taken from *International Financial Statistics* (IMF) in the section "Cost and Price Comparisons in Manufacturing." We use the relative normalized unit labor costs (line 65umc 110). The U.S. price of oil comes from *DRI*. It is the average U.S. refiners' acquisition price of imported crude oil. U.S. nominal interest rates are taken from the *Bank of Canada Review* (table 20). We use the rates on 3-month commercial paper. Data for the U.S. GNE deflator are taken from *DRI*. The LDCs' terms of trade are taken from *International Financial Statistics* (IMF). They are defined as the ratio of export prices of non-oil LDCs (line 74d, 201) to import prices of non-oil LDCs (line 75d, 201).

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ECONOMIES OF SCALE AND THE LOAD FACTOR IN ELECTRICITY GENERATION

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Abstract—We present three alternative ways of introducing the load factor into a typical cost function employed in the analysis of electricity generation. Relying on the translog flexible functional form, these three alternatives are estimated with plant data under two maintained hypotheses, an exogenous load factor and an endogenous load factor. Under the former maintained hypothesis, a non-nested hypothesis test favors the alternative which treats the load factor as a state of nature variable; under the latter maintained hypothesis, statistical and economic considerations favor the alternative which treats the load factor as a proxy for speed variations. In both cases, however, there are significant economies of scale over observed output levels despite the introduction of the load factor, which contradicts a claim made in the earlier literature.

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Introduction

While there are many empirical studies of the electricity generation process, there is a limited consensus on what these results show.[1] In particular, Stewart (1979) questions the validity of the neoclassical production approach applied to plants as units. He stresses that this approach ignores the nature of output in an electricity generation plant. Namely, it must be thought of in at least two dimensions, i.e., an instantaneous rate, which measures potential, and a cumulative rate of actual output, which depends on utilization over a period. Employing an engineering production approach, Stewart comes to the surprising conclusion that plant size has no statistically significant effect on costs when the load factor is included in the analysis. This result contradicts, of course, a common empirical finding in the literature. In this paper, we address the issue raised by Stewart's work from the perspective of models based on

[1] For example, Cowing and Smith (1978) and Joskow and Schmalensee (1983).

the neoclassical production approach that incorporate the load factor into the analysis.

In section I, we present the models that underlie the empirical analysis. Each of those models represents an alternative way of introducing the load factor into the standard cost function, which is the basis for many empirical analyses of electricity generation, for example, Christensen and Greene (1976). The first model (I) views the load factor as a "state of nature" variable; hence, it treats this variable in the manner suggested by Stevenson (1980). The second one (II) views the load factor as a measure of duration; hence, it includes this variable in the manner suggested by the model developed in Betancourt (1986). The third one (III) views the load factor as a measure of intensity or speed, and it is explicitly developed in section I.

Subsequently, in section II, we discuss the empirical implementation of the three models in terms of the translog flexible functional form. Since these three models are not special cases of each other, we present a non-nested hypothesis test to be used, among other criteria, in comparing the performance of these models. This section also contains a brief description of the plant data base, which was originally developed by McFadden (1978), and a discussion of the issues raised by the potential endogeneity of the load factor.

The paper concludes with a presentation of the results of estimating these models under the maintained hypothesis that the load factor is exogenous as well as under the maintained hypothesis that it is endogenous. Under the former hypothesis, all three variants of the neoclassical model show that plant size is an important determinant of costs even though the load factor is included in the empirical analysis. Under the latter hypothesis, however, this result holds for models II and III but not for model I. Since there is a powerful reason for rejecting model I under this endogeneity hypothesis, our results definitely contradict the claim by Stewart (1979, p. 564) that "the major cost reduction at the unit level comes from increases in the plant utilization factor not from increases in the size of the unit..."

I. Alternative Ways of Modelling the Load Factor

In this section, we set out three plausible theoretical models that can be used to incorporate the load factor into the structure of production in electricity generation.

Model I postulates a neoclassical cost function of the following form:

$$TC = C(P_1, P_2, P_3, Y; LF) \tag{1}$$

where TC are total costs, P_1 represents the cost of capital (K), P_2 is the price of labor services (L), P_3 denotes the price of fuel (F), Y is the actual level of output and LF is the load factor. Note that, by definition, $LF = Y/X$ where X is the capacity level of output, usually measured in terms of the rated capacity of the equipment. The function C has the following properties: continuous, increasing, linear homogeneous and concave in prices, increasing in output and twice differentiable with respect to all its arguments. In this model, which is adapted from Stevenson's (1980) work on technical change, the load factor is viewed as a state of nature variable, i.e., an exogenous characteristic of the production process which will also affect costs.

In model II the load factor is viewed as a measure of the duration of operations. Hence, the specification of the cost function is derived from the model in Betancourt (1986), i.e.,

$$TC = LF\, C(P_1^*, P_2, P_3, X), \tag{2}$$

where the asterisk is used to indicate the dependence of the capital service price on the load factor or the duration of operations. Specifically, $P_1^* = P_1/LF$, and it represents the price of capital services per unit of duration of operations.[2] The duration of operations could also affect the price of labor services or fuels if variations in duration require the firm to pay shift differentials or experience variations in start up or shutting down costs.[3] The use of X in the cost function represents the selection of full capacity utilization, $LF = 1$, as the reference unit of duration of operations.

Finally, in model III the load factor is viewed as a measure of the speed or intensity of operations for any given level of duration. In this case, the cost function can be specified as

$$TC = C(P_1^*, P_2, P_3, Y). \tag{3}$$

It is derived as follows: The producer wants to minimize costs that are given by

$$TC = P_1 K + P_2 L + P_3 F, \tag{4}$$

subject to the production function constraint which is given by

$$Y = F[(LF)K, L, F]. \tag{5}$$

For ease of exposition (4) is written as

$$\begin{aligned} TC &= (P_1/LF)(LF)K + P_2 L + P_3 F \\ &= P_1^*(LF)K + P_2 L + P_3 F. \end{aligned} \tag{4'}$$

The result of the optimization problem is a set of conditional input demand functions for capital services $[(LF)K]$, labor services (L) and fuel flow (F). These functions are substituted into (4′) in order to generate the general form described in (3).

[2] The numerator in this expression can also depend on the load factor if one allows for increasing wear and tear as the duration of operation increases.

[3] Since the dependence of P_2 and P_3 on duration is implicit, rather than explicit as in the case of capital services, it will be dealt with upon estimation.

II. Empirical Implementation

For purposes of comparability, the empirical implementation is undertaken using the same flexible functional form to approximate the respective cost functions, i.e., the function $C(\ )$ in all three models. The translog form is employed because of its simplicity and its prior use in the literature on electricity generation. The properties of symmetry of the second partials and linear homogeneity in prices are imposed on the translog for each model prior to estimation. The properties of monotonicity and concavity are used to check the consistency of the models with the data.

Estimation of translog cost functions in similar contexts is usually undertaken jointly with the estimation of the input shares in costs, which are obtained from the cost function via Shephard's Lemma, e.g., Christensen and Greene (1976). We follow the literature in this procedure as well as in deleting one of the share equations and using the iterative version of Zellner's (1962) seemingly unrelated regression technique to obtain maximum likelihood estimates. Table 1 presents the translog equation estimated for the cost function in each model with symmetry and linear homogeneity imposed. The corresponding equations for the shares of capital and fuel are omitted to conserve space; the share equation for labor was the one deleted in the estimation.

A comparison of the three equations in table 1 suffices to realize that none of the three models is a special case of the other ones. Therefore, a comparison of their performance is undertaken by relying on the recent literature on non-nested hypotheses tests. As MacKinnon (1983) points out, for example, one can write the null hypothesis in a multivariate context as

$$\mathrm{H}_0: Y_{it} = f_{it}(\beta) + u_{it}^0,\ u_t^0 \sim (0, \Omega_0), \tag{6}$$

and the alternative as

$$\mathrm{H}_1: Y_{it} = g_{it}(\gamma) + u_{it}^1,\ u_t^1 \sim (0, \Omega_1), \tag{7}$$

where i indexes the equations, t indexes the observations and Ω_j is the contemporaneous covariance matrix for the error terms corresponding to hypothesis j. One can then construct a compound model as follows:

$$\mathrm{H}_C: Y_{it} = (1-\theta) f_{it}(\beta) + \theta g_{it}(\hat{\gamma}_1) + u_{it}, \tag{8}$$

where $g_{it}(\hat{\gamma}_1)$ is obtained by replacing γ with its maximum likelihood estimator under the alternative hypothesis. If H_0 is true u_t should have covariance matrix Ω_0. Joint estimation of θ and β can be undertaken using (8), which provides the basis for the so-called J-test. Recently, Smith and Maddala (1983) have generalized this procedure to the testing of multiple models through the use of a compound model of the form

$$\mathrm{H}_C: Y_{it} = \left(1 - \sum_{j=1}^{m} \theta_j\right) f_{it}(\beta) + \sum_{j=1}^{m} \theta_j g_{it}(\hat{\gamma}_j) + u_{it}^*. \tag{9}$$

TABLE 1.—TRANSLOG COST FUNCTIONS

Model I

$$\begin{aligned}\ln TC = {} & \alpha_0 + \alpha_1(\ln P_1) + \alpha_3(\ln P_3) + \alpha_Y \ln Y + \alpha_{LF}\ln LF \\ & + \gamma_{11}(\ln P_1)^2/2 + \gamma_{33}(\ln P_3)^2/2 + \gamma_{YY}(\ln Y)^2/2 \\ & + \gamma_{LF}(\ln LF)^2/2 + \gamma_{13}[(\ln P_1)(\ln P_3)] \\ & + \gamma_{Y1}\ln Y(\ln P_1) + \gamma_{Y3}\ln Y(\ln P_3) \\ & + \gamma_{YLF}\ln Y \ln LF \\ & + \gamma_{LF1}\ln LF(\ln P_1) + \gamma_{LF3}\ln LF(\ln P_3)\end{aligned}$$

Model II

$$\begin{aligned}\ln TC = {} & \alpha_0 + \alpha_1(\ln P_1^*) + \alpha_3(\ln P_3) \\ & + \alpha_X \ln X + \alpha_{LF}\ln LF \\ & + \gamma_{11}(\ln P_1^*)^2/2 + \gamma_{33}(\ln P_3)^2/2 \\ & + \gamma_{XX}(\ln X)^2/2 \\ & + \gamma_{13}[(\ln P_1^*)(\ln P_3)] + \gamma_{X1}\ln X(\ln P_1^*) \\ & + \gamma_{X3}\ln X(\ln P_3)\end{aligned}$$

Model III

$$\begin{aligned}\ln TC = {} & \alpha_0 + \alpha_1(\ln P_1^*) + \alpha_3(\ln P_3) + \alpha_Y \ln Y \\ & + \gamma_{11}(\ln P_1^*)^2/2 + \gamma_{33}(\ln P_3)^2/2 \\ & + \gamma_{YY}(\ln Y)^2/2 \\ & + \gamma_{13}[(\ln P_1^*)(\ln P_3)] + \gamma_{Y1}\ln Y(\ln P_1^*) \\ & + \gamma_{Y3}\ln Y(\ln P_3)\end{aligned}$$

Note: All terms in the table involving prices and total costs are normalized by the price of labor services, e.g., $(\ln P_1)$ is actually $(\ln P_1 - \ln P_2)$ and so on. This notational device is not used anywhere else in the paper.

Following this procedure, the J-test of one model against the other two will be conducted through the use of a likelihood ratio test.

Our data base consists of a set of 36 plants constructed during the 1957–61 period.[4] This data base was originally developed by McFadden (1978) from information available in two publications of the Federal Power Commission, *Steam-Electric Plant Construction Cost and Annual Production Expenses (Annual Supplements)* and *Statistics of Electric Utilities in the United States*. In general, we have followed McFadden's procedures in constructing the data used in the analysis from the original sources; in particular, we have constructed the variables relevant for the analysis of costs during the plant's first full year of operation. This approach is useful because, as McFadden argues, it provides a means of controlling for technological change. The procedure for construction of the variables employed in the empirical analysis (corresponding to Y, X, LF, P_1, P_1^*, P_2, P_3, TC, P_1K/TC, P_1^*K/TC, P_2L/TC, and P_3F/TC) on the basis of these two sources is described in detail in an appendix available upon request. Finally, use of a data base established in the literature ensures that our results are not due to creative data construction.

Past research on electricity generation has frequently assumed the exogeneity of the load factor. This is true of studies relying on plant data; for example, McFadden (1978, pp. 98–99) discusses this assumption and Stewart (1979) implicitly adopts this assumption in

[4] Of the 36 plants 28 are single generating unit plants and at least 6 of the remaining plants contain two generating units.

TABLE 2.—ESTIMATED VALUES OF SCE

	1	2	3	4	5
Plant Size[a]	0.217	0.600	0.779	1.089	3.562
A. Exogenous Load Factor					
Model I	1.28	0.929	0.834	0.734	0.326
	(136.08)	(1063.30)	(1373.5)	(730.42)	(94.77)
Model II	0.501	0.398	0.302	0.251	−0.040
	(54.33)	(162.39)	(140.83)	(116.65)	(−10.19)
Model III	0.486	0.272	0.213	0.155	−0.097
	(39.46)	(80.75)	(62.98)	(25.58)	(−3.84)
B. Endogenous Load Factor					
Model I	0.528	0.591	0.488	0.452	0.234
	(0.129)	(0.191)	(0.159)	(0.152)	(0.073)
Model II	0.493	0.370	0.259	0.198	−0.141
	(3.30)	(2.77)	(2.04)	(1.52)	(−1.00)
Model III	0.494	0.275	0.215	0.156	−0.099
	(8.23)	(5.50)	(4.30)	(3.12)	(−1.41)
Load Factor[b]	0.33	0.65	0.63	0.71	0.90

Note: t-ratios are given in parentheses below the coefficient estimates.
[a]Plant size measured in millions of megawatt hours of actual output for the median plant in each group.
[b]Value of the load factor for the median plant in each group.

his regression analysis (table 6). It is also true of studies relying on firm level data (Stevenson, 1980). The previous discussion has been based on this maintained hypothesis. As one referee pointed out, however, the assumption is more defensible at the firm level than at the plant level, since a utility can meet its exogenous level of demand through reallocation of output among plants as well as purchases from networks. Therefore, all three models were also estimated under the maintained hypothesis of endogeneity of the load factor using nonlinear three stage least squares as the estimation method.

Before presenting the results, it is worth noting how the hypothesis of endogeneity affects the interpretation of these models. In model I, endogeneity implies that the state of nature interpretation is not valid. Instead, the model can be viewed as one in which the load factor is treated as an input and the cost function is specified in terms of the input because its price is not observable. Incidentally, this interpretation was developed for the standard case by Halvorson and Smith (1986). Model II is consistent with the load factor being either exogenous or endogenous. In the latter case, however, the possible dependence of the prices of labor services and fuel on the load factor suggests the specification of one or both of these variables as endogenous upon estimation. Finally, model III implies that capital services are chosen through the choice of the load factor (capital stock) given the level of the capital stock (load factor) when the load factor is endogenous (exogenous). Hence, it is also consistent with both maintained hypotheses.

III. Results

In presenting the results, we shall concentrate on those relevant for the main substantive issue raised in the paper: namely, the nature of economies of scale in the presence of the load factor. For details on other characteristics of the production structure, the reader is referred to the earlier version of the paper.

Economies of scale are defined in terms of the cost function as

$$SCE = 1 - \partial \ln C / \partial \ln Y \tag{10a}$$

$$SCE = 1 - \partial \ln C / \partial \ln X, \tag{10b}$$

where the first equation is applied to models I and III and the second equation is applied to model II. In either case, this concept measures the proportional increase in cost as a result of a proportional increase in size, measured by actual output in (10a) and capacity output in (10b). This measure results in positive numbers for economies of scale and in negative numbers for diseconomies of scale. The top half of table 2 presents estimates of SCE for each model under the maintained hypothesis of exogeneity of the load factor. These estimates are presented for the median observation in a five group partition of the sample, where the groups are ordered by the level of actual output.

For all three models, the null hypothesis of constant returns to scale is rejected at levels of significance far beyond the 1% level. The pattern of the estimates is one of a consistent decline in economies of scale as the size of the median plant increases between groups. The levels of these estimates differ substantially between model I and the other two models. That is, in model I the median plant in the largest size group is still experiencing substantial economies of scale whereas in the other two models the same plant is already experiencing diseconomies of scale. The non-nested hypothesis test based on the likelihood ratio yielded the following observed values of the test statistic when each of the three models is employed as the null hypothesis in a comparison of (6) and (9): model I, 6.96; model II, 39.88; model III, 20.74. At the 1% level of significance, the critical value of a χ^2 with two degrees of freedom is 9.21. Therefore, model I cannot be rejected while the other two are rejected at the 1% level of significance and, thus, the data support model I over the other two models.

In evaluating these results, mention must be made of the consistency of each model with the data as revealed, for example, by satisfaction of the concavity conditions for each observation. Five observations failed to satisfy the concavity conditions in all three models. In addition, another observation failed to satisfy concavity in models II and III and a seventh failed the test in model III. None of these 7 observations coincided with the median ones for which estimates are reported in table 2. Incidentally, monotonicity was satisfied by all observations in each model.

One way of interpreting these results is that the translog model is an approximation and as long as the concavity conditions are satisfied at the relevant points in the data set, for example at the median plants selected in table 2 or at the mean of the values of the variables, one should not be concerned about the existence of violations of concavity for some other observations. Alternatively, one can argue that violations of these conditions for a "substantial" part of the sample is a serious problem which casts doubts on the validity of the maintained hypotheses underlying the model in which these violations occur. In our particular case, it is more fruitful to discuss the results of relaxing the assumption that the load factor is exogenous than to attempt a reconciliation of these two interpretations.

The bottom part of table 2 presents the results of estimating the same three models under the assumption of endogeneity of the load factor.[5] In model I the null hypothesis of constant returns to scale cannot be rejected at any reasonable level of significance. The reason is that in model I, when the load factor is endogenous, neither plant size nor the load factor itself are statistically significant explanatory variables (for example at the 5% level) in the determination of total costs. By contrast, the hypothesis of constant returns to scale is rejected at the 5% level in (three) four of the five groups in model (II) III. Moreover, the pattern of the estimates for these two models is similar to the ones obtained under the exogeneity hypothesis. Indeed, the actual estimates for model III under both maintained hypotheses differ only in the third decimal.

These results lead to the rejection of model I as an empirically viable mechanism for incorporating plant size and the load factor simultaneously into the analysis of electricity generation under the hypothesis of endogeneity of the load factor. Consequently, they also cast doubt on the underlying economic model which is based on treating the load factor as another input in a restricted cost function.

A check of the concavity conditions for each of the models based on the NL3SLS estimates was performed. The condition is violated by two observations in model I, twenty-seven in model II and four in model III. These results lead to the rejection of model II because of its violation of the cost minimization hypothesis for a substantial number of observations. In addition, model II can be subjected to another consistency test. This model implies that the coefficient of the load factor in table 1 (α_{LF}) should equal unity. This hypothesis was rejected by the data at the 1% level of significance under the exogeneity hypothesis but not under the endogeneity hypothesis.[6]

To sum up, plant size has a statistically significant role in the determination of costs when the load factor is included in the analysis. Under the hypothesis of endogeneity of the load factor, however, model III, or the speed interpretation, is the most sensible way of including the load factor in the analysis on the basis of both statistical and economic considerations.

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[5] In models I and III P_1, P_2, P_3 and Y are exogenous; hence, the endogeneity of the load factor ($LF = Y/X$) is induced by the endogeneity of design output (X). By contrast, in model II P_1, P_2 and X are exogenous; hence, the endogeneity of the load factor is induced by the endogeneity of actual output (Y). NL3SLS was implemented by using as instruments the terms of a second degree polynomial in the logarithms of the exogenous variables for each model.

[6] Incidentally, the results for model II in table 2 do not incorporate this restriction ($\alpha_{LF} = 1$). Estimation of model II under the hypothesis of endogeneity of the load factor and imposing the restriction ($\alpha_{LF} = 1$) leads to violations of the concavity conditions for twenty-four observations and to estimates that are very close to the ones in table 2.

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ECONOMIES OF SCALE VERSUS TECHNOLOGICAL CHANGE IN THE NATURAL GAS TRANSMISSION INDUSTRY

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Abstract—Although most productivity studies to date assume away scale economies, those studies which both incorporate scale economies and provide a relevant breakdown find that scale economies dominate technological change as a primary determinant of productivity growth. This paper measures the determinants of productivity growth in the U.S. Natural Gas Transmission industry. The technology was estimated using a translog approximation. Results of the econometric model were then utilized to measure the determinants of productivity growth in this industry. It was found that technological change explained certainly as much and often more of productivity growth as did scale economies.

I. Introduction

With some notable exceptions, much of the empirical literature on productivity measurement disregards the potential importance of scale economies. Rather, constant returns to scale are assumed to prevail and technological change is presumed to be the principal determinant of productivity growth.[1] On the other hand, those studies which do incorporate scale economies in their productivity measures find that scale economies explain at least twice as much productivity growth as does technological change.[2] This result, as others have pointed out (e.g., Denny, Fuss and Waverman (1981), p. 206), is rather surprising.

The purpose of this paper is to analyse the components of productivity growth in the U.S. Natural Gas Transmission Industry for the years 1953 to 1979 with a view to determining the relative importance of technological change versus scale economies in generating growth. There are a number of reasons for singling out this industry. First, the technology per se is rather simple. There is only a single output, namely natural gas, which is transported by pipeline over long distances. Second, an upper bound estimate for the scale economy factor, namely 2.07, has been derived from laboratory experiments on pipeline design and used in engineering production function analyses of the gas transmission industry.[3] It is only an upper bound since actual pipeline design is adversely affected, among other things, by terrain and population density factors which are absent in the laboratory.[4] Third, the available data

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[1] See, for example, May and Denny (1979) and Gollop and Jorgenson (1980).

[2] See, for example, Denny, Fuss and Waverman (1981), Nadiri and Schankerman (1981), and Chan and Mountain (1981).

[3] See, for example, Robinson (1972). Previous economic studies which have made use of such laboratory production functions include Callen (1978) and Aivazian and Callen (1981).

[4] Safety requirements, for example, necessitate that pipelines near populated areas have a thicker wall thereby reducing scale economies.