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THE USE OF CROSS-SECTIONAL ESTIMATES OF PROFIT FUNCTIONS FOR TESTS OF RELATIVE EFFICIENCY: A CRITICAL REVIEW*

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As a result of the development of duality theory, cost functions and profit functions have been used in a number of recent studies of production. In a number of applications, it is substantially easier to obtain good estimates of these functions than of the traditional production function and duality theory shows that any 'well-behaved' cost or profit function corresponds to a neoclassical production function.

One major difficulty with this approach may arise when cross-sectional data are used. The main arguments of cost and profit functions are price variables and these are not likely to vary greatly between firms at a single point in time. This point has been raised before (see, for example, Christensen and Greene 1976; Varian 1978, p. 124). However, many authors have either ignored it or have adopted procedures which create spurious variations in measured prices. The most serious of these spurious variations result from the incorporation of quality differences in 'price' variables.

The problems associated with cross-sectional estimation of cost and profit functions are illustrated using the methodology developed by Lau and Yotopoulos (1971, 1972) and Yotopoulos and Lau (1973) for testing relative efficiency. This is an inherently cross-sectional approach and in most applications wages are the main price variable used. In this note, the effects of using quality variables in place of prices are explored and the results derived by Lau and Yotopoulos and other writers using their methodology are examined critically.

Sources of Measured Price Variation

In a perfectly competitive market, all firms at a given time and place face the same vector of prices. In cross-sectional studies, differences in time are excluded so that variation in the vector of actual prices faced by firms can come only from differences in location or from violations of the competitive assumptions.

If differences of location are an important source of price variation, then there does not appear to be any theoretical difficulty in estimating cost and profit functions, or in testing for efficiency differences between groups of firms. However, this is not likely to be the case in many applications, particularly agricultural applications. If farms are so widely dispersed that they face significantly different factor and output prices, then climatic variations are likely to invalidate the vital assumption that all firms have the same technology (apart from possible multiplicative differences in 'technical' efficiency). This is particularly

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important in cross-sectional studies where climatic experience in a single year must be assumed to be constant across firms, unless it is explicitly incorporated during estimation.

Price variations due to violations of the competitive assumptions will permit estimation of a cost function only if all firms act as price takers. This condition might be satisfied if some firms received input subsidies which were independent of their own actions. For example, access to cheap credit might be confined to members of specific ethnic or religious groups. The condition would not be satisfied if firms had monopoly power or if they faced factor prices set by a discriminating monopolist. Price differences due to information costs create significant difficulties both in estimation and interpretation, and are not likely to provide a generally satisfactory basis for work of this type.

Another possible cause of price variation is variation in the opportunity cost of operator and family labour. For a farm on which hired labour is not used, it is possible that the opportunity cost of operator and family labour at the margin is less than the market wage. In such a situation, however, the competitive assumption that all factors are in infinitely elastic supply to the firm at the current price (in this case, the marginal opportunity cost) is violated. It is, therefore, impossible to use the standard methods of duality theory to determine profit maximising input levels.

Thus, the user of cross-sectional price data faces a dilemma. If perfectly competitive conditions prevail, the methodology is theoretically valid, but the absence of any variation in the data will make estimation of profit functions impossible. If, on the other hand, there is variation in prices, it is unlikely that profit functions can validly be derived from competitive assumptions.

The most important danger in this area is, however, that, even though the actual vectors of prices faced by firms are equal, the measured prices are not. This problem is particularly likely to arise when information about prices is obtained from the individual firm, rather than from observations of the relevant market. One possible source of error is simple measurement error. Operators may not state the prices actually paid. This danger is more acute when the price variable is a notional one, for example, the inputed wage of family labour. Note that this error will also affect measures of total cost and profit.

A more complex problem is that of quality differences. Most inputs are heterogeneous in quality and different firms will employ inputs of various quality levels in different proportions. If the price paid by the firm is measured by dividing total payments to that factor by the aggregate of total services of each factor, without regard to quality differences, then price and quality differences will be confounded. In particular, in the competitive case when all firms face the same vector of prices, the 'price' measured by this procedure will in fact be an index of factor quality.

Factor Share Equations and the Profit Maximisation Assumption

One solution to the dilemma outlined above is the use of additional information from factor share equations. The assumption of profit maximisation may be adopted as a maintained hypothesis, and used to derive equality constraints between the share equations and the profit function.

Indeed, in the extreme case of perfect competition with no measured price variations, all of the information about the coefficients of the profit function must be derived from the share equations.

There are a number of problems associated with this approach. First, there is good evidence that, in a risky environment, farmers will prefer to maximise expected utility rather than expected profit. Second, the validity of the constraints depends on the correctness of the assumed functional form for the profit function. It is very difficult to test this assumption if the coefficients of the unrestricted profit function are insignificant or, worse still, spurious.

Despite these problems, the use of profit maximisation as a maintained hypothesis in the estimation of cost and profit functions seems to be a defensible procedure. However, very serious problems arise with procedures such as that of Lau and Yotopoulos, in which an attempt is made to test the hypothesis.

In the absence of data problems, it is relatively straightforward to specify appropriate tests for the profit maximisation hypothesis. Given an appropriate profit function and associated factor share equations estimated by generalised least squares, profit maximisation implies a series of cross-equation linear restrictions. The main small-sample tests which may be used are Wald statistics (generalised t-tests) likelihood ratio statistics and Lagrange multiplier statistics. Although these test statistics have asymptotically identical chi-square distributions, Berndt and Savin (1977) and Breusch (1979) have shown that, for well-behaved problems, the Wald statistic is always larger than the likelihood ratio, which in turn is larger than the Lagrange multiplier test, while Theil (1971, pp. 402-3) has shown that the use of the asymptotic χ^2 is more likely to lead to the rejection of the null hypothesis than any of these.

Theil argued for the use of the more 'cautious' small-sample measures which reduce the risk of Type I error. However, for the problem at hand the main danger is that of a Type II error. Because of the lack of variation in price data derived from competitive markets, it is likely that the null hypothesis of profit maximisation will be accepted even when it is not valid. A failure to reject the null hypothesis should not be taken as strong evidence in favour of profit maximisation since an alternative null hypothesis might have performed equally well.

The problems associated with using cross-sectional price data to test the profit maximisation assumption may be illustrated using the Lau-Yotopoulos test of relative efficiency. In the following section, the test is described and a number of studies which use it are examined.

The Lau-Yotopoulos Relative Efficiency Test

Lau and Yotopoulos (1971, 1972) assumed Cobb-Douglas technology. Firms are assumed to have fixed endowments of inputs $Z_1, \ldots, Z_m$ and to be free to choose variable inputs $X_1, \ldots, X_n$. The variable inputs are not necessarily chosen to maximise profits. Rather thay are set by equating marginal cost of factor i to $1/k_i$ times its marginal value product. Firms are described as allocatively efficient if all the k_i are equal to unity.

Firms are divided into two groups and it is assumed that technical efficiency may differ between the two groups by a multiplicative factor. The ratios k_i are assumed to be constant within groups but may differ between groups. This yields an estimated unit output price profit function

$$\ln\Pi = A + \delta_1 D_1 + \sum_{i=1}^{n} a_i \ln C_i + \sum_{j=1}^{m} \beta_j \ln Z_j \tag{1}$$

and factor share equations

$$-C_i X_i/\Pi = \alpha_{li} D_1 + \alpha_{2i}(1-D_1)\ i = 1, 2, \ldots n \tag{2}$$

where Π is unit output price profit (total revenue less total variable costs, normalised by the price of output), C_i are factor prices, normalised by output price, and D_1 is a dummy variable taking the value one for farms in group 1 and zero for those in group 2.

The hypothesis of equal economic efficiency between groups implies $\delta_1 = 0$. The hypothesis that group 1 is profit maximising implies $\alpha_{li} = \alpha_{2i}$ for $i = 1, 2, \ldots, n$ and similarly for group 2. The hypothesis of equal price efficiency between the two groups implies $\alpha_{1i} = a_{2i}$ for $i = 1, 2, \ldots, n$.

Equation (1) and the factor share equations are estimated using Zellner's seemingly unrelated regression method (Zellner 1962). Any of the restrictions $\delta_1 = 0$, $\alpha_{li} = \alpha_i$ or $\alpha_{li} = \alpha_{2i}$ may be imposed and tested. Elasticity estimates may be derived with or without these restrictions.

This methodology, with minor variations, was followed by Sidhu (1974), O'Connor and Hammonds (1975), Trosper (1978), Khan and Maki (1979) and Sidhu and Baanante (1979). Sidhu and Baanante[1] presented restricted estimates only, while O'Connor and Hammonds presented unrestricted estimates only. (It is not clear whether the latter were derived using ordinary least squares or Zellner's method.)

In most of the studies, labour is the only variable factor. Sidhu and Baanante also treated chemical fertiliser and irrigation water as variable inputs while O'Connor and Hammonds, in their study of meat marketing, treated wholesale meat as variable.

All of the measured prices derived for these variable factors appear to be open to the strictures laid down above. Sidhu and Trosper used the average wage rate reported by the firm for hired labour and imputed the same rate to family labour (with adjustments for child and family labour). Khan and Maki and Lau and Yotopoulos used different imputed rates and formed the wage rate as a weighted average of the imputed family rate and the reported rate for hired labour. O'Connor and Hammonds and Sidhu and Baanante did not state how their wage rates or other variable factor prices were derived. All of the above authors, except Lau and Yotopoulos, compared firms within a single region, and it is therefore unlikely that transport costs would cause significant variations in factor prices. In the case of Lau and Yotopoulos (1971), data from five Indian states were used, but state dummy variables were introduced to account for differences in output prices between states. These would also account for any interstate differences in wage rates. Lau and Yotopoulos were the only authors to include their data. These data show little variation in wage rates within states, although differences between states are substantial.

[1] Following the results of Sidhu (1974), Sidhu and Baanante took equal price efficiency between small and large farms as a maintained hypothesis. Thus, the results they referred to as 'unrestricted' correspond to those described here as having one restriction imposed.

TABLE 1

Parameter Estimates in Unit Output Price Profit Model[a]

Author(s)	Variable	Column 1		Column 2		Column 3		Single equation ordinary least squares	
		Zellner's method							
		Unrestricted		One restriction[b]		Two restrictions[c]			
Yotopoulos and Lau (1973)	Labour	−0.90	(−1.0)	−0.98	(−1.1)	−1.14	(−3.1)	−2.14	(−2.1)
Khan and Maki									
Punjab	Labour	−0.03	(−0.2)	−0.11	(−0.8)	−0.15	(−28.4)	−0.32	(−2.0)
Sind	Labour	−0.13	(−1.1)	−0.15	(−1.2)	−0.39	(−4.9)	−0.28	(−2.0)
Trosper	Labour	−0.004	(−0.01)	na		−0.03	(−6.8)	0.08	(0.3)
O'Connor and Hammonds	Wholesale meat	−4.34	(−1.0)	na		na		na	
	Labour	−4.01	(−1.3)	na		na		na	
Sidhu									
1967-68	Labour	0.26	(1.9)	0.26	(1.9)	−0.24	(−7.2)	0.11	(0.7)
1968-69	Labour	0.02	(0.2)	0.02	(0.2)	−0.38	(−9.3)	0.51	(−2.4)
1969-70	Labour	−0.06	(−0.5)	−0.6	(−0.5)	−0.25	(−3.1)	−0.28	(−2.2)
1970-71	Labour	−0.18	(−1.0)	−0.18	(−1.0)	−0.25	(−10.0)	−0.48	(−2.5)
1967-68 to 1970-71	Labour	−0.08	(−1.2)	−0.08	(−1.2)	−0.28	(−6.6)	−0.24	(−3.1)
1969-70	Labour	−0.06	(−0.6)	−0.06	(−0.6)	−0.25	(−3.1)	−0.29	(−2.3)
Sidhu and Baanante	Labour	na		−1.10	(−0.5)	−0.44	(−15.7)	−0.72	(−2.5)
	Fertiliser	na		−0.42	(−0.8)	−0.16	(−20.1)	−0.87	(−1.1)
	Irrigation	na		−0.42	(−0.4)	−0.11	(−12.0)	−0.03	(−0.4)

[a] The numbers in parentheses are t-statistics.
[b] One restriction: $\alpha_{1i} = \alpha_{2i}$.
[c] Two restrictions: $\alpha_{1i} = \alpha_i$; $\alpha_{2i} = \alpha_i$.

A particularly interesting feature of the dummy variable approach used by Lau and Yotopoulos (1971, p. 105) is the assumption that 'the prices of outputs differ only across states'. The reasoning behind this assumption could equally well be applied to input prices but this would render a test of allocative efficiency impossible.

The crucial test of the usefulness of cross-sectional price variables is, of course, their performance in estimation. The results from previous studies are not encouraging. Of twelve coefficients estimated for price variables (by unrestricted Zellner's) in eleven equations, only two were significantly different from zero at the 20 per cent level and one of those had the wrong sign. (See Table 1, column 1.) This performance is marginally worse than that which would be expected by chance. From the results reported by Sidhu and Baanante, it appears that all three of their coefficients would be insignificant in the absence of restrictions.

Given these results, severe doubt is cast on the appropriateness of the test of absolute price efficiency, that is, the profit maximising assumption. As Trosper (1978, p. 512) states 'The test of absolute price efficiency is not very convincing because the standard error of the labour coefficient is large in the profit equation'. This large standard error means that the null hypothesis is very unlikely to be rejected, even when the measured price variable is completely specious, and indeed only Sidhu, and Khan and Maki rejected the hypothesis for any group of farmers. In most of these cases of rejection, the coefficient was the wrong sign or else very near zero in value.

The bias toward acceptance of the null hypothesis is shown by the fact that Khan and Maki (for Punjab) and O'Connor and Hammonds both rejected the hypothesis of equal price efficiency, while accepting that both groups were absolutely price efficient. This is due to the fact that the standard errors on the coefficients in the factor share equations are much lower than those of the corresponding price variable in the profit function.

Even apart from this extreme case, the difficulties with the labour coefficient in the profit equation render the test of equal relative price efficiency extremely dubious. It is clear that, if two producers have identical Cobb-Douglas technology except for a multiplicative parameter but have significantly different labour shares, they cannot both be price efficient. However, it is impossible to tell which is more price efficient without an accurate estimate of the labour coefficient.

One reason why the problems associated with these price variables have not attracted greater attention may be the fact that the unrestricted Zellner's estimates (which are central to the methodology) are usually presented along with several other estimates, which frequently have highly significant coefficients on prices (see Table 1). Typically, the additional estimates are derived from single equation ordinary least squares or from Zellner's method with equality restrictions on the various price coefficients a_i, α_{1i} and α_{2i}. The reason why the price coefficients become significant in the second case is clear. The coefficients in the factor demand equations (which are simply estimates of the mean factor share) are naturally highly significant and this desirable property carries over to the coefficient in the profit function once equality constraints between the two equations are imposed (see Table 1, column 3).

The improvement obtained using single equation ordinary least squares is mainly due to the fact that the coefficient estimates are more strongly negative than those obtained using unrestricted Zellner's. As shown in Table 1, this occurs in every case where both estimates are reported.

This could be explained if variation in measured prices were due to reporting or measurement errors. An excessively high reported price will impart a corresponding upward bias to the measured factor share and a downward bias to measured profits. This means that the error in the factor share equation will be closely correlated with the factor price variable in the profit function equation and hence with the error term in this equation when the coefficient of 'price' is near zero. The use of the seemingly unrelated regression estimator will lead to this correlation being taken into account, and hence to the magnitude of the price coefficient being reduced.

A similar argument may apply when quality differences are the main source of variation in measured prices. This will yield an excessively high ordinary least squares coefficient estimate whenever factor quality is correlated with factor shares.

The behaviour of the coefficients in the share equations is also of some interest. The imposition of the within-equation constraint $\alpha_{1i} = \alpha_{2i}$ is essentially equivalent to pooling two samples to obtain an estimate of the mean. Thus it is not surprising that the significance of the coefficients is increased. The subsequent imposition of the cross-equation constraint provides a more critical test. If the input price variable in the profit function contains useful information, and the allocative efficiency hypothesis is correct, the imposition of the constraint should improve the significance of the share equation coefficient. On the other hand, the analysis presented here suggests that the imposition of the constraint simply transfers significance from the share equation to the profit function. This process will not improve the significance of the share equation coefficient(s), and may reduce it.

The latter view is supported by results given in Table 2. The imposition of the within-equation restrictions was accepted only by Yotopoulos and Lau (1973) and Sidhu (1974). Both show improvements in significance. By contrast, the subsequent imposition of the cross-equation constraints has no major effects on significance. The t-statistics are improved slightly in two cases, and worsened slightly in six. This is consistent with the hypothesis that no additional information is gained from the profit function.

Thus, both the way in which 'price' variables are derived and the results of econometric estimation support the hypothesis that any measured variations in these variables is spurious. It is noteworthy that most of the studies discussed above deal with traditional agriculture. It might be suggested that, in this case, markets would be less well developed and prices more susceptible to variation than in modern agriculture or other industries. Given the disappointing results presented here, it is likely that application of the profit function methodology to cross-sectional data for modern industries is likely to create serious problems.

This conclusion does not apply to efficiency tests alone. For example, a number of the papers discussed above contain estimates of input elasticities. If there is only one variable factor, the elasticity is given by

TABLE 2

Coefficients in Factor Share Equations[a]

Author(s)	Variable	Estimates using Zellner's method							
		Unrestricted[b]				One restriction[c]		Two restrictions[d]	
Yotopoulos and Lau (1973)	Labour	−1.82	(−2.9)	−0.66	(−1.2)	−1.21	(−3.1)	−1.14	(−3.1)
Khan and Maki									
Punjab	Labour	−0.26	(−29.8)	−0.11	(−21.5)	−0.15	(−28.4)	−0.15	(28.4)
Sind	Labour	−1.40	(−7.0)	−0.28	(−2.4)	−0.56	(−5.3)	−0.39	(−4.9)
Sidhu									
1967-68	Labour	−0.29	(−7.2)	−0.22	(−2.9)	−0.27	(−7.8)	−0.24	(−7.2)
1968-69	Labour	−0.43	(−7.3)	−0.41	(−6.2)	−0.42	(−9.8)	−0.38	(−9.3)
1969-70	Labour	−0.44	(−2.2)	−0.50	(−3.2)	−0.48	(−3.9)	−0.25	(−3.1)
1970-71	Labour	−0.26	(−5.7)	−0.25	(−5.0)	−0.26	(−10.4)	−0.26	(−10.4)
1967-68 to 1970-71	Labour	−0.35	(−4.9)	−0.41	(−5.3)	−0.38	(−7.3)	−0.28	(−6.7)
Sidhu and Baanante	Labour	na		na		−0.46	(−15.9)	−0.44	(−15.7)
	Fertiliser	na		na		−0.16	(−18.2)	−0.16	(−20.0)
	Irrigation	na		na		−0.11	(−11.4)	−0.11	(−12.0)

[a] The numbers in parentheses are t-statistics.
[b] The coefficient for small farms is given first.
[c] One restriction: $\alpha_{1i} = \alpha_{2i}$.
[d] Two restrictions: $\alpha_{1i} = \alpha_i$; $\alpha_{2i} = \alpha_i$.

$-\alpha_1/(1-a_1)$. The variance of this estimate is given by $\sigma_\alpha^2/(1-\alpha_1)^2$, where σ_α^2 is the variance of α_1. (See Kendall and Stuart 1969, p. 232.)

Since the distribution of the elasticity estimates will not, in general, be normal, t-statistics cannot be presented. It is possible, however, to make some observations on the relative standard errors. (These will differ by a factor of $1-\alpha$ from those for the input price coefficients from which they are derived.) They are quite large when unrestricted estimates are used. The imposition of the constraint $\alpha_{1i}=\alpha_{2i}$ (that is, equal allocative efficiency) on the cost share equation does not have much effect. On the other hand, after the imposition of the cross-equation restriction $\alpha_1=\alpha_{1i}=\alpha_{2i}$, the largest relative standard error is 0.15. Note that in a number of cases the restricted and unrestricted elasticity estimates differ by as much as a factor of four. These results are presented in Table 3.

In general, the restricted estimates are fairly similar to those which would be obtained using estimates of mean factor shares and Cobb-Douglas technology. However, the methodology used here tends to endow them with a largely spurious air of sophistication.

Alternative Approaches

The difficulties associated with the uniformity of competitively determined prices vitiate the usefulness of the Lau-Yotopoulos methodology in many situations. It is, then, incumbent on users of this approach to establish that there are special factors operating in a particular situation, which justify the use of the Lau-Yotopoulos methodology.

This justification may be undertaken in three stages. First, it is necessary to establish that there is, indeed, substantial variation in the observed price data. This may be done by reporting coefficients of variations. It is also desirable to report parameter estimates and significance levels for the profit function in the absence of restrictions. Second, the obvious sources of data variation, such as measurement error and quality differences, must be considered. If these are large enough to account for the observed variation, or a large part of it, then the approach should be avoided. Third, it is necessary to explain the observed variation in a manner consistent with the hypothesis that all farmers are price takers.

The third of these tasks is the most difficult. As noted above, in a sample drawn from geographically disparate, but climatically similar farms, transport costs could yield an adequate explanation. Alternatively, in some cases, government policy might create differential prices. However, neither of these conditions is likely to be satisfied regularly. A third possibility relates to variation in prices over time.

The closest approach to meeting these requirements has been made by Flinn, Kalirajan and Castillo (1982) in a study of Filipino rice growers. They reported coefficients of variation for their price data, which range from 9 per cent to 28 per cent. They argued that quality differences were unlikely to be important sources of measured price variation for fertiliser and rice which are fairly standard products. (However, it should be noted that these two prices had the lowest coefficients of variation, 9 per cent and 11 per cent, respectively.) The observed price differences were attributed to transport cost, and variations in dealer prices (which were not explained).

Given the difficulties associated with price variation, there will be

TABLE 3

Indirect Estimates of Input Elasticities for Labour[a]

Author(s)	Restrictions None	One[b]	Two[c]
Yotopoulos and Lau (1973)	0.47 (0.25)	0.49 (0.23)	0.53 (0.08)
Khan and Maki			
Punjab	0.03 (0.12)	0.10 (0.12)	0.13 (0.00004)
Sind	0.12 (0.10)	0.13 (0.09)	0.28 (0.04)
Trosper	0.0004 (0.25)	na	0.03 (0.0004)
Sidhu			
1967-68	−0.36 (0.25)	−0.36 (0.25)	0.20 (0.02)
1968-69	−0.02 (0.15)	−0.02 (0.15)	0.28 (0.02)
1969-70	0.05 (0.09)	0.05 (0.09)	0.20 (0.05)
1970-71	0.15 (0.13)	0.15 (0.13)	0.20 (0.02)
1967-68 to 1970-71	0.08 (0.06)	0.085 (0.06)	0.22 (0.03)
1969-70	0.06 (0.10)	0.06 (0.10)	0.20 (0.05)

[a] The numbers in parentheses are standard errors.
[b] $\alpha_{1i} = \alpha_{2i}$.
[c] $\alpha_{1i} = \alpha_{2i} = \alpha_i$.

many cases in which the Lau-Yotopoulos technique is not appropriate. In such cases, the most promising approach appears to be the traditional one of estimating a production function, with the use of additional information from factor share equations. The main objection to this approach has been the claim of simultaneous equation bias associated with the use of input levels as explanatory variables. An associated objection is the 'regression fallacy' pointed out by Stigler (1952) in relation to size economies, where random variations in demand levels could yield spurious estimates of size economies.

However, as has been pointed by Zellner, Kmenta and Dreze (1966) and Vlastuin, Lawrence and Quiggin (1982), among others, these objections are not really pertinent in an agricultural context, where output variations are mainly due to climatic factors which become apparent after input decisions have been made. In this case, single equation ordinary least squares estimates of production function parameters are unbiased, provided price and climatic fluctuations are independent of each other (see Zellner et al. 1966, pp. 790-1).

The use of production functions is not entirely free of problems. In particular, if production technology is given by

$$\ln Y = \Sigma_i \alpha_i \ln X_i + \delta_i \tag{3}$$

where δ_i is distributed as $N(0,1)$, then $E[Y]$ is not equal to exp ($E[\ln Y]$). As a result, there is a bias in ordinary least squares share equation estimates of the α_i which creates difficulties in tests of efficiency.

Where possible, it is desirable to use information from both production functions, and profit functions, and compare these with the relationships suggested by duality theory. While very little work has been done on econometric testing of dual relationships, it is desirable to compare direct and indirect estimate of elasticities, as has been done by Flinn et al. (1982).

Concluding Comments

The use of increasingly sophisticated econometric techniques, with maintained hypotheses suggested by economic theory, offers the potential of more efficient estimation of economic models. However, there is frequently a price to be paid, in terms of reduced robustness to violations of the maintained hypotheses.

This problem becomes particularly critical in the case of cross-sectional studies using price data. In general, sufficient data for estimation will be available only if the maintained hypothesis of perfectly competitive markets is violated. This problem may be overcome by adopting an additional maintained hypothesis that all firms choose input levels so as to maximise profits, but it is not generally possible to provide a valid test of this hypothesis using price data.

If the Lau-Yotopoulos test of relative efficiency is to be used, it must be established that there is sufficient variation in prices to provide a meaningful test of efficiency, and that this variation arises from factors which do not violate the assumption that all firms are price takers. If these conditions are not fulfilled, alternative approaches, such as those based on production functions, must be used.

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