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Energy substitutability in transition agriculture: estimates and implications for Hungary

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Abstract

Subsidised energy prices in pre-transition Hungary had led to excessive energy intensity in the agricultural sector. Transition has resulted in steep input price increases. In this study, Allen and Morishima elasticities of substitution are estimated to study the effects of these price changes on energy use, chemical input use, capital formation and employment. Panel data methods, Generalised Method of Moments (GMM) and instrument exogeneity tests are used to specify and estimate technology and substitution elasticities. Results indicate that indirect price policy may be effective in controlling energy consumption. The sustained increases in energy and chemical input prices have worked together to restrict energy and chemical input use, and the substitutability between energy, capital and labour has prevented the capital shrinkage and agricultural unemployment situations from being worse. The Hungarian push towards lower energy intensity may be best pursued through sustained energy price increases rather than capital subsidies.

JEL classification: Q4; P2

Keywords: Hungarian agriculture; Energy use; Transition; Allen and Morishima elasticities of substitution; Generalised Method of Moments

1. Introduction

The degree of substitutability between energy and other inputs in the production process has been an important area of research for applied economists, both in manufacturing and in agriculture. Much of the prior work in this area has related to the market economies of the developed world, with the motivation arising initially from steep oil price rises, and later from environmental and non-renewable resource conservation concerns. Thus, while input substitution can be viewed as a purely technical relationship, it also has obviously important policy ramifications. To take an example, if capital and energy are strongly complementary, persistent and large energy price increases will discourage capital formation and thus limit the long-run growth of the agricultural sector.

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Although compelling research has been conducted on the evolution of commodity prices, outputs and institutional features in the agricultural sectors of Central and East European transition economies for example, Macours and Swinnen (2000), Mathijs and Swinnen (1998), research focused explicitly on energy use and substitutability has been very limited. However, it is a particularly important research question for these countries because they have been significant consumers of energy, and have been witnessing significant changes in the structure of input subsidies and relative input prices in the process of transition. Further changes are anticipated as some of the countries prepare for accession to the EU.

This paper attempts to bridge this gap in knowledge by estimating elasticities of substitution between energy and other inputs in Hungarian agriculture, and analysing the economic and policy implications. Allen as well as Morishima elasticities of substitution are estimated and used jointly to derive implications. In contrast to most of the previous works that have used either cross-sectional or aggregate time-series data in estimating substitution elasticities, we use farm-level panel data, thereby accounting for both year- and farm-specific effects. An effort is made to be as careful as possible in the econometric model specification and estimation—in particular, a pure production function approach is taken to avoid behavioural misspecification, a possibly significant danger when dealing with newly established or incomplete market economies. Recently developed econometric methods involving panel methods, Generalised Method of Moments (GMM) estimation, and a series of instrument exogeneity tests are used to specify and estimate the model. Implications of the substitution elasticities for energy use by the Hungarian agricultural sector are derived. In particular, the estimates enable us to analyse how evolving energy prices in transitional Hungary may have affected capital formation, employment and chemical input use, and whether further energy conservation may be best accomplished by direct price policy or by tax credits on energy-saving capital.

Section 2 surveys energy use in Hungary and Section 3 reviews the theory of input substitution. Section 4 outlines the estimation techniques and the data are briefly discussed in Section 5. The results are reported in Section 6 and the implications are discussed in Section 7. The conclusion briefly summarises the findings.

2. Energy use in Hungarian agriculture

As is well known, Hungary had a predominantly command economy structure until the late 1980s, when the transition to a market economy gathered momentum. Energy prices for all sectors under the command structure had been heavily subsidised, resulting in low energy efficiency and large energy intensity of production. In 1993, the energy intensity per unit of GDP in Hungary was 0.73 tonnes of oil equivalent per 1000 US$, compared to an average of 0.40 for the OECD (OECD, 1995). Although the bulk of Hungary's total energy consumption and inefficiency is attributable to the residential and heavy industrial sectors, problems have been observed in the agricultural sector as well. For example, Marrese (1993, p. 141) noted that "... Hungarian agriculture became too energy-intensive as Hungarian farmers responded to artificially low prices".

Energy prices have rapidly increased since the initiation of reforms in 1989, almost quintupling by the end of the first decade of reform. While the primary motivation for the removal of subsidies has been the commitment to market orientation, a secondary motive has been the concern with excessive energy intensity. This motive has its origins not only in environmental and non-renewable energy conservation concerns, but also in the dependence on foreign imports. In Hungary, only about 54% of primary energy is produced domestically, and there is a heavy dependence on imports, particularly from the Russian Federation. In April 1993, the Hungarian Parliament adopted a new energy policy committing itself to achieving greater energy efficiency as a matter of urgency (OECD, 1995). The principal tools to attaining this have included not only price policy (subsidy removal), but also the promotion of more energy-saving capital.

However, it is not just the energy prices that have changed during the transition process in the agricultural sector. Subsidy and price support removals have been across the board, resulting in soaring input prices and fluctuating output prices. These relative price changes raise several questions regarding energy use in the post-transition period. Have the price increases for the other inputs discouraged or encouraged energy use? Have the escalating energy prices promoted or inhibited capital formation and agricultural employment? Are the twin policies of energy use restriction and promotion of energy-saving capital best accomplished by continued energy price increases or via subsidies on energy-saving capital?

Information on elasticities of substitution can help answer these important questions. However, no estimates of substitution elasticities between energy and other inputs in post-transition Hungarian agriculture are available, an important reason for this being the paucity of appropriate data. Accurate estimation of desired substitution elasticities without introducing substantial aggregation problems requires farm-level data, including information on energy use. Within the context of farm level data, the econometric importance attached to controlling for time and farm-specific effects points to the need for panel data. There is a marked lack of availability of such data for post-transition Hungary. While the European Commission's FADN data collection for Hungary commenced in 1998, it will be some years before a panel long enough in the time dimension is available.

We fortunately have access to a panel dataset of 117 Hungarian farms for the years 1985–1991, and in this paper, we use this pre-transition/early-transition dataset to estimate substitution elasticities. We then use these estimates to derive implications for the immediate late-transition/post-transition period, assuming implicitly that technical change has not distorted the picture. While this is a strong assumption, the view taken here is that in the face of a lack of alternative information and data this is a valid and useful exercise. As Chambers (1991, p. 118) notes in the context of the gap between data needs and availability for applied production analysis, "The solution is either to simplify the problem or give up since the data is insufficient to permit further analysis. [...] The second course seems rather nihilistic and dooms the economist to either purely theoretical problems or analysis of extremely simple problems. Usually, one is better off with some information, no matter how imperfect, than with no information".

Obtaining substitution elasticity estimates through direct technology estimation instead of dual methods also enables us to avoid pitfalls associated with behavioural specification in a transition period. This point is elaborated later.

3. Measurement of input substitution

The production technology is defined by a production function given by:

$$Y = f(X_1, X_2, \dots, X_n) \tag{1}$$

where Y is the output, and $X_1, \dots, X_n$ are the inputs. The direct (or Hicksian) elasticity of substitution between any two inputs i and j attempts to measure the curvature of the isoquant at a given point in (i, j) space, given by

$$\sigma_{i,j} = \left(\frac{\mathrm{d}(X_j/X_i)}{X_j/X_i}\right)\left(\frac{f_i/f_j}{\mathrm{d}(f_i/f_j)}\right) \tag{2}$$

As shown in Eq. (2), the direct elasticity of substitution measures the proportional change in two inputs in response to a change in their marginal rate of technical substitution, holding all other inputs constant. However, in a production function with more than two inputs, it is hard to imagine all other inputs staying fixed while only two are allowed to change. Thus, Allen (1938) presented an alternate general expression for the elasticity of substitution that does not make such an assumption. The Allen elasticity of substitution (AES) is defined as:

$$\sigma^{\mathrm{A}}_{i,j} = \frac{\sum_i X_i f_i}{X_i X_j}\frac{F_{ji}}{F} \tag{3}$$

where F is the determinant of the bordered Hessian for the production function, and F_{ji} is the cofactor associated with f_{ij} in the bordered Hessian. The AES is symmetric, so that $\sigma^{\mathrm{A}}_{i,j} = \sigma^{\mathrm{A}}_{j,i}$. When $\sigma^{\mathrm{A}}_{i,j} > 0$, inputs i and j are substitutes, and when $\sigma^{\mathrm{A}}_{i,j} < 0$, they are complements.

The AES continues to be the most commonly reported measure of input substitution, even in recent studies of input substitution (e.g. Hisnanick and Kyer, 1995). However, it is long known to have several shortcomings and severe doubts have been cast upon its usefulness by critics. Blackorby and Russell (1989) demonstrated that $\sigma^{\mathrm{A}}_{i,j}$ is just the cross-price elasticity of input demand divided by the cost share of the jth input. While the cross-price elasticity is an intuitive and policy-relevant quantity, the AES disguises this quantity by dividing by a cost-share. Most empirical elasticity of substitution studies estimate cost functions, and can therefore directly provide measures of cross-price elasticities. The AES thus does not provide any additional information, and indeed, may confound useful information. Blackorby and Russell also showed that in the multiple input case the AES does not measure curvature in the natural way that Eq. (2) does in the two input case.

The Morishima elasticity of substitution (MES), first put forth by Morishima (1967) and brought to prominence by Blackorby and Russell (1981, 1989), is an alternative that provides information on isoquant curvature and retains the intuition associated with Eq. (2) in a setting with more than two inputs. In terms of the production function, the MES can be written as

$$\sigma^{\mathrm{M}}_{i,j} = \frac{f_i}{X_i}\frac{F_{ij}}{F} - \frac{f_j}{X_j}\frac{F_{ji}}{F} \qquad (4)$$

The MES is not a symmetric measure, i.e. $\sigma^{\mathrm{M}}_{i,j} \neq \sigma^{\mathrm{M}}_{j,i}$. Inputs that are classified as AES compliments may be classified as MES substitutes, although inputs classified as AES substitutes will continue to be classified as MES substitutes. An elaborate discussion of other aspects of the conceptual desirability of MES compared to AES, and on the properties of the MES can be found in Chambers (1991). It is noted that there are additional reasons to prefer the MES—prior empirical applications have found that for inputs that have relatively small cost shares, relatively small changes in the use of the input can induce large changes in AES estimates (Thompson and Taylor, 1995). Additionally, the MES has been found to be significantly more robust to levels of data aggregation than the AES (Nguyen and Streitwieser, 1997).

4. Econometric approach

4.1. Choice of methods

Elasticities of substitution are most easily estimated using the dual cost function, assuming cost-minimising equilibrium in competitive markets. Detailed and reliable input price data are not available to us, however, and this option is directly ruled out. The alternative then is direct estimation of the primal equation, Eq. (1). However, direct, single-equation, least squares estimation of production functions can suffer from simultaneous equation bias. As has been known since the work of Marschak and Andrews (1944), empirical estimation has to contend with the reality that input levels are endogenous variables chosen by producers. The endogeneity creates a correlation between the input levels and the production error term, resulting in inconsistent least squares estimates. Although an escape route is available by appealing to the Zellner–Kemnta–Dreze proposition that simultaneity bias is not a problem if producers maximise *expected* profits (Zellner et al., 1966), researchers can seldom use this defence without misgivings (Griliches and Mairesse, 1998).

One option, following some studies that have estimated substitution effects using the production function (Nguyen and Streitwieser, 1997; Humphrey and Moroney, 1975), is to invoke the assumptions of competitive markets with cost-minimising production, setting up a system of simultaneous equations. Specifically, under the assumption that each input is used such that its value marginal product equals the input price, cost share equations can be derived and estimated as a system of equations. This approach, apart from addressing the simultaneity bias concern, can also provide more efficient estimates than single-equation production function estimation provided the assumptions of cost minimisation and competitive markets are valid.

However, imposing these behavioural assumptions for Hungary in the pre-transition/early-transition period may be unwarranted. The sample covers 1985–1991. In the pre-transition period, i.e. 1985–1988, the farms operated within a command economy and could not be characterised well by standard neo-classical behavioural models. Price reforms were initiated in 1989, and proceeded at great speed thereafter. Thus, the later years of our sample can be said to represent a time of substantial uncertainty regarding input and output prices in Hungarian agriculture. The potential for deviation from informed cost-minimising behaviour (behavioural 'errors') was severe, which brings into question the validity and usefulness of the share equations. Therefore, single-equation estimation of the production function is the most reasonable choice, and accounting for simultaneity becomes important in this case. Even in the simple case with no producer-specific heterogeneity, there is cause to suspect that the inputs in the production function will be correlated with the error term, as elaborated above. This correlation becomes even more likely when unobserved heterogeneity among producers is present (for example, land quality).

Our approach to tackling this is to use panel methods to eliminate heterogeneity effects and to use instrumental variable (IV) techniques. There is an additional problem—the only IVs available in our dataset are the past, present and future values of the (endogenous) input quantities. In other words, the IVs may only be pre-determined and not necessarily exogenous. In the last decade, a series of papers have been published that confront the problems of simultaneity and unobservable correlated effects in panel data IV estimation by devising systematic specification tests (Mairesse and Hall, 1996; Arellano and Bond, 1991; Keane and Runkle, 1992). The approach is to set up a succession of models that are appropriate under various assumptions regarding instrument validity, and then using GMM methods to choose between them. This is the general approach taken in this paper, and is sketched briefly in the following section.[1]

4.2. Aspects of panel data and methods

Suppose the production model in Eq. (1) could be written in log-linear form. For convenience, we specify only a single input in this exposition. Then, the panel model can be written as

$$y_{it} = \beta x_{it} + \alpha_i + \lambda_t + e_{it} \quad (5)$$

where y is the logarithm of output, x the logarithm of input, i indexes farms, and t represents years. λ_t is the time-specific effect that affects all producers equally in a year (such as a drought). α_i represents the farm-specific 'fixed effect' that is presumed correlated with the x_{it} (such as soil quality). The time-effects are the easiest to handle since the removal of the year means (across farms, within a specific year) from the variables gets rid of λ_t. We continue our discussion by assuming that year effects have been already removed in this fashion from Eq. (5).

An accepted way of removing the farm-specific effects is first-differencing.[2]

$$y_{it} - y_{it-1} = \beta(x_{it} - x_{it-1}) + e_{it} - e_{it-1} \quad (6)$$

Estimation of Eq. (6) by least squares will provide consistent estimates provided e_t is a pure, unanticipated shock that does not transmit to x_t and x_{t-1}. If there is cause to suspect 'residual simultaneity', the first-difference equation has to be instrumented. In an agricultural production context, one circumstance leading to lingering simultaneity problems (even after fixed effects are eliminated) is when some inputs are applied on the basis of sequential decision making. For example, pesticides may be applied sequentially in a given year in response to unravelling information about pest infestations. This would cause pesticide inputs in year t to be correlated with the production error term in year t.

Where values of the input variables in the dataset are the only instruments available, different years in the panel will have different numbers of instruments available to them. For example, in a 7 years panel such as ours, if strong exogeneity is not valid and only current and past years' input variables can be used as instruments, then the last period (first difference between years 7 and 6) will have more instruments available to it than the period preceding it and so on. GMM enables such IV estimation with differing numbers of instruments available to different years.

Denoting first-differencing by Δ for notational convenience, we can write Eq. (6) as

$$\Delta y_{it} = \beta \Delta x_{it} + \Delta e_{it} \quad (7)$$

If the input from a time period s is a valid instrument for Δx_{it}, the orthogonality condition can be written as

$$E(\Delta e_{it} x_{is}) = 0 \quad (8)$$

where $\Delta e_{it} = \Delta y_{it} - \beta \Delta x_{it}$. In practice, the number of orthogonality conditions depends upon the number of inputs x and the number of years from which the inputs can be regarded as valid instruments. As noted earlier, this implies that different years will have different numbers of instruments available. For example, under strong exogeneity, Δe_{it} is uncorrelated with inputs from all years in the sample, and we could write

[1] An excellent discussion can be found in Griliches and Mairesse (1998).

[2] The 'within' transformation is another option, but has been criticised by Chamberlain (1982), among others.

$E(\Delta e_{it} x_{is}) = 0$ $(s = 1, \dots, t)$. Under weak exogeneity, Δe_{it} is uncorrelated only with past and present values of inputs, and we write

$$E(\Delta e_{it} x_{is}) = 0, \qquad s = 1, \dots, t$$

4.3. GMM estimation and specification tests

We do not discuss GMM estimation methods in any detail here,[3] restricting ourselves to a brief sketch. In general, a vector of parameters are to be estimated from moment conditions given by

$$E(\varepsilon_i(\beta) \otimes v_i) = 0 \tag{9}$$

where $\varepsilon_i(\beta)$ is a set of disturbance terms, such as the Δe_{it}, $\otimes$ is the Kronecker product, and v_i is a set of valid instruments. These moment conditions have sample analogues given by

$$h(\beta) \equiv \frac{1}{N}\sum_{i=1}^{N} h_i(\beta) = \frac{1}{N}\sum_{i=1}^{N} \varepsilon_i(\beta) \otimes v_i \tag{10}$$

The sample moments are combined and minimised with respect to β, i.e. choose β to minimise the quadratic form

$$A(\beta) \equiv h'(\beta)\Sigma h(\beta)$$

where Σ is a positive definite weighting matrix. The choice of a consistent estimator of the inverse of the covariance matrix of $h(\beta)$ as Σ yields consistent and asymptotically efficient estimates of β. GMM estimation using panel data requires weaker distributional assumptions than some alternate methods, and is robust to serial correlation and heteroskedasticity in the panel. Where there are overidentifying restrictions (the number of moment conditions exceed the number of parameters to be estimated), a chi-square test of the overidentifying restrictions can be used, and interpreted as a check on the internal consistency of the instruments. The minimised value of the GMM criterion function evaluated at the estimated parameters (Hansen's J-statistic) is distributed asymptotically as a chi-square variable with degrees of freedom equal to the number of instruments less the number of parameters, and forms the basis of the overidentifying restrictions test.

Even after first-differencing, as in Eq. (7), there can be residual simultaneity in the model. Previous experience with panel data on firms (Mairesse and Hall, 1996; Hall, 1992) suggests that such simultaneity may render invalid assumptions of strong exogeneity (instruments from all years valid for equations of all years), weak exogeneity (all past and present instruments valid), lag1 instruments (all instruments up to $(t-1)$ valid for year t equation), lag2 instruments (all instruments up to $(t-2)$ valid for year t equation), etc., in decreasing order of likelihood.[4] Instead of assuming the validity of a certain set of instruments, it is wiser to test to determine the instrument lag structure that is valid for the data and model being used. This can be done by a set of nested chi-square tests. We start with the strongest assumption, strong exogeneity (all instruments for all years) and test it against the next strongest, i.e. weak exogeneity (lag0+ instruments). Then weak exogeneity is tested against the slightly weaker assumption of lag1+ instruments, and finally lag1+ instruments is tested against the model with the weakest assumptions in the set we consider, that with lag2+ instruments. In the case of each pair, this is done by calculating the difference between the two Hansen's J-statistics, producing a statistic that is asymptotically distributed as chi-squared, with degrees of freedom equal to the number of extra moments implied by the additional instruments. The final choice of instrument set is made on the basis of the nested tests.

4.4. The translog production function

Since the ultimate objective of this study is to present information on substitution possibilities between inputs, a flexible functional form is required that does not constrain the elasticities of substitution to be constant. The translog is an obvious choice, since apart from enabling elasticities of substitution to vary across input levels, the translog is interpretable as a second order approximation to an underlying unobservable production function. Additionally, it is

[3] Mairesse and Hall (1996) provide a good discussion, with direct relevance to production function estimation and simultaneity matters. Caselli et al. (1996) also provide an accessible exposition of these methods in an empirical growth setting.

[4] Here, 'equation for year t' means the equation first-difference between $(t-1)$ and t.

log-linear, enabling us to avoid non-linear estimation. The general form of the translog for the 4 input case is given by

$$\ln Y = \ln A_0 + \sum_{k=1}^{4} A_k \ln(X_k) + \frac{1}{2}\sum_{k=1}^{4}\sum_{l=1}^{4}\beta_{kl}\ln(X_k)\ln(X_l) \qquad (11)$$

where k and l index the inputs. Of course, the translog is not necessarily globally well-behaved in practice. One regularity check that is easy with the translog is that of monotonicity, i.e. $\partial Y/\partial X > 0$ at chosen input levels, which can be accomplished by checking that the output elasticity, $\partial \ln(Y)/\partial \ln(X)$ (in this case linear in parameters), is positive.

5. Data

The data are a balanced panel of 117 farms for the period 1985–1991, compiled by the Hungarian Ministry of Finance. Data are available on aggregate farm output (value of gross output) and four inputs: materials (fertilisers, pesticides and seed), energy (fuels and electricity), labour and capital. Apart from labour, on which information on both wage bill and number of employees is available, all the inputs are in current value terms in the original dataset. Although the capital input should ideally be expressed as a flow, this proved to be impossible in this case because of a lack of adequate information on depreciation and interest rate variables. Therefore, the capital variable used is simply the gross value of fixed assets. Information on land size was not available, and therefore it must be assumed that their effects are incorporated in the unobservable farm-specific effects.

Unfortunately, price information was not available at the farm level. Given the nature of price movements during the period, it was especially important to account for inflation across time. Deflator series were constructed from price indices reported in the Hungarian national statistics (Statistical Yearbook of Hungary, 1992), and the output and input variables were converted into constant values. Some details about the data, including summary statistics, are presented in the appendix. For further details regarding the dataset and its transformation, see Piesse (1999), where the data have been used in the analysis of efficiency issues in Hungarian agriculture.

6. Estimation and results

The translog production Eq. (11) was estimated using GMM methods after removing time effects and first-differencing to remove fixed-effects. Of the 7 years of data available, five were used in specifying first-difference production function equations. Thus, there were four first-difference production function equations, pertaining to 1991–1090, 1990–1989, 1989–1988, and 1988–1987. Since the weakest in our suite of models is the one with instruments from $(t-2)$ and backwards, data for 1985 and 1986 were used purely to provide instruments. In all cases, symmetry ($\beta_{kl} = \beta_{lk}$) was imposed prior to estimation.

Four models, 'strong exogeneity' (all instruments for all years), 'weak exogeneity' (all past and present values of inputs as instruments), 'lag1+' (inputs upto $(t-1)$ as instruments for t) and 'lag2+' (inputs upto $(t-2)$ as instruments for t) were estimated and nested chi-squared tests (based on differences between J-statistics from each pair of models) used for selection. For these tests to be valid, it is necessary that the estimated covariance matrix of orthogonality conditions $\hat{\Omega}$ from the model with the weakest assumption be used as the weighting matrix for all other models. Accordingly, we have used $\hat{\Omega}$ estimated from the lag2+ model to weight the other models. The results are presented in Fig. 1.

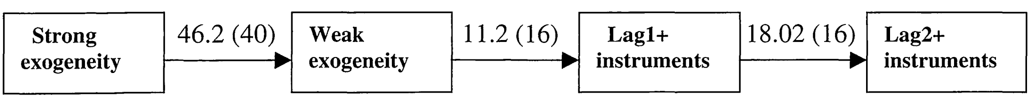

Fig. 1. Testing for exogeneity, 117 Hungarian farms, 1985–1991 (pairs of chi-square tests, degrees of freedom in brackets; H_0: additional moments implied by stronger exogeneity assumptions are valid).

Table 1
Translog production function GMM estimates

Regressor	Parameter estimate	S.E.	P-value
Fixed effects—strong exogeneity[a]			
Materials	0.1718	0.1457	0.238
Energy	−0.1806	0.2175	0.406
Labour	1.8261	0.2378	0.000
Capital	0.0527	0.1561	0.736
Materials × materials	0.3421	0.0376	0.000
Energy × energy	0.1157	0.0316	0.000
Labour × labour	−0.1092	0.0627	0.082
Capital × capital	−0.0524	0.0244	0.032
Materials × energy	−0.0978	0.0310	0.002
Materials × labour	−0.0690	0.0318	0.030
Materials × capital	−0.1690	0.0196	0.000
Energy × capital	0.0126	0.0296	0.000
Energy × labour	−0.1034	0.0303	0.001
Labour × capital	0.1250	0.0357	0.000
Fixed effects—lag2+ instruments[b]			
Materials	−0.0177	1.0437	0.986
Energy	−0.9013	1.4264	0.527
Labour	3.0897	1.0776	0.004
Capital	0.4726	0.7218	0.513
Materials × materials	0.3819	0.2127	0.073
Energy × energy	−0.0661	0.1954	0.735
Labour × labour	−0.1490	0.2789	0.593
Capital × capital	−0.2756	0.1208	0.023
Materials × energy	−0.1066	0.1935	0.582
Materials × labour	−0.1150	0.1529	0.452
Materials × capital	−0.1475	0.0923	0.110
Energy × capital	0.3527	0.1904	0.064
Energy × labour	−0.1052	0.1601	0.511
Labour × capital	0.0982	0.1370	0.473

[a] Hansen's J-statistic test of overidentifying restrictions: 87.8700 ($P = 0.759$).
[b] Hansen's J-statistic test of overidentifying restrictions: 12.3511 ($P = 0.989$).

As seen from Fig. 1, at the 1% level, none of the null hypotheses are rejected and we are able to choose the model with the strongest assumption, that of strong exogeneity. This implies that there is no evidence to indicate residual contemporaneous or lagged correlation problems, and that the first-difference errors are indeed 'pure', untransmitted errors.[5] Estimates of the translog production parameters for this chosen model are presented in the upper half of Table 1. For comparison, estimates from the model with the weakest assumption, i.e. lag2+ instruments are reported in the lower half of the table.

Table 2
Output elasticities at sample means

Elasticity for	Estimate	Error	t-statistic	P-value
Materials	0.490880	0.012712	38.6149	0.000
Energy	0.105291	0.015369	6.85096	0.000
Labour	0.367008	0.012923	28.3993	0.000
Capital	0.054279	0.010341	5.24897	0.000
Sum	1.01745			

Table 1 shows the substantial improvement in the precision with which parameters are estimated when the chosen strong exogeneity model is used instead of the lag2+ model. Ten of the fourteen parameters are significant at the 5% level. Hansen's J-statistic has a P-value of 0.759, indicating that the overidentifying restrictions are jointly orthogonal to the error term.

Having chosen the strong exogeneity model, output elasticities are calculated at the overall sample means of the inputs. Table 2 shows that the elasticities are positive, reasonable and significant at the 1% level. The elasticities sum to 1.017, which suggests constant returns to scale. Since the translog production function does not impose monotonicity, it is essential to check monotonicity at individual data points. We do this by checking whether the output elasticities are positive at each sample observation. Monotonicity with respect to materials is satisfied at 99% of the observations. The corresponding numbers for energy, labour and capital inputs are 89, 99 and 72%, respectively, the last number reflecting perhaps that the stock representation of capital input is less than ideal.

7. Substitution elasticities and their implications

We proceed to use expressions (3) and (4) to calculate the Allen and Morishima elasticities, respectively, at the sample mean values. As noted before, the elasticities of substitution have been estimated using a single-equation production function approach because the data run over the early transition period, for which the scope for behavioural misspecification is high. However, the interesting implications of substitution elasticities flow from using them to infer

[5] Although one typically expects some residual correlation problems in firm data, Mairesse and Hall (1996) also found that the data accepted strong exogeneity in their analysis of French manufacturing firms over 1981–1985.

the responses of input quantities to price changes in other inputs, i.e. by assuming cost minimisation in the late/post-transition period and thereby the equality between MRTS and input price ratios. The rapid adjustment of the transition countries towards market orientation, as noted by several authors, makes such interpretations plausible. For example, Trzeciak-Duval (1999) notes that farmers in the transition countries appeared to be significantly more responsive in the second half of the 1990s to relative price signals in making production decisions. Another assumption made in using the substitution elasticities thus is that the technology estimates from these pre/early-transition data accurately represent technology in the late/post-transition period. This is doubtless a strong assumption, particularly given technical change possibilities. However, given the importance of the substitutability issue, and the paucity of data and research of this kind for Hungary, we make the assumption with the view that the insights generated may be useful despite any shortcomings. Another aspect that could limit the validity of such analysis in transitional countries is that these periods are often marked by input unavailability at the farm level. If input shortages exist, there can only be limited substitution between inputs at the margin. However, the literature indicates that the availability of commercial inputs in Hungary was not one of the many significant problems that the agricultural sector faced during transition. Hungary possessed a better-supplied input market than most transitional countries, and was able to lean on imports when the input industry started to downsize. In the case of labour and (physical) capital inputs, shortages were not an issue since numerous farms went out of business, creating a surplus situation.

Another point to be noted, before analysing the implications of the estimates, is that clear forecasts of input use patterns cannot be made in an atmosphere where everything, from input and output prices, to institutional features of the agricultural sector, are changing. We can only discuss *tendencies* for input substitution at the margin (output held constant) implied by the technology estimates. In Hungary, the price of practically every input has increased dramatically through the 1990s, while output prices first fell sharply and later stabilised. Naturally, the result has been a substantial initial contraction of the sector, characterised by sharp reductions in the use of all inputs, followed by a moderate recovery.

One of the criticisms of the AES has been that it does not provide any information in addition to the cross-price elasticity of input demands, and in fact is not as easily interpretable as the cross-price elasticity (Blackorby and Russell, 1989). However, since a cost function is not estimated, and an estimate of the cross-price elasticity is not directly available to us, the AES estimates do provide useful information. With four inputs, there are six elasticities, which are reported in Table 3(A). Note that standard errors are not calculated for the substitution elasticities because the computations are excessively complicated. Standard errors for AES are relatively easy to compute when estimation is accomplished via the cost function, as noted by Binswanger (1974). The AES in terms of the translog production function is an exceedingly complex, non-linear function of several parameters, making calculation of standard errors infeasible (Humphrey and Moroney, 1975). Indeed, these calculations are typically not attempted (see Bregman et al. (1995) and Nguyen and Streitwieser (1997)). The case of the MES is even more complicated.

As classified by Mundlak (1968), the AES is a one-price one-factor elasticity of substitution, i.e. it

Table 3
Elasticities of substitution

Pairwise substitution elasticities	
(A) Allen elasticities	
Materials/energy	−3.59
Materials/labour	2.06
Materials/capital	0.15
Energy/labour	4.58
Energy/capital	7.95
Capital/labour	−2.10
(B) Morishima elasticities	
Materials/energy	−0.02
Materials/labour	2.10
Materials/capital	0.14
Energy/labour	3.01
Energy/capital	0.56
Capital/labour	0.59
Energy/materials	−1.35
Labour/materials	1.37
Capital/materials	0.45
Labour/energy	0.82
Capital/energy	1.16
Labour/capital	0.02

Table 4
Price indices for selected agricultural inputs in Hungary[a]

Input	1990	1992	1994	1996	1998
Fuels	100	173	224	359	494
Fertilisers	100	146	199	373	421
Pesticides	100	160	202	324	389

[a] *Source*: Hungarian Central Statistical Office (1999).

Table 5
Quantity (consumption) indices for selected agricultural inputs in Hungary[a]

Input	1990	1992	1994	1996	1998
Fuels	100	65	62	60	62
Fertilisers	100	31	50	48	59
Pesticides	100	64	69	54	49

[a] *Source*: Hungarian Central Statistical Office (1999).

captures the effect of the change in one input's price upon the use of another input. The first point to note is that the AES estimates involving energy are the largest in the set, implying that energy use in Hungarian agriculture is particularly sensitive to the prices of other inputs. All AES estimates involving energy are in excess of 3.5 in absolute value, while the largest non-energy AES, for capital and labour, is 2.10 in absolute value. While this sensitivity cautions us that changes in the prices of non-energy inputs can have strong unintended effects upon energy use, it also points to the feasibility of expanding the set of policy options beyond the usual own price control strategy. It appears that there exists some potential to control energy use in Hungarian agriculture via indirect means.

The AES between energy and materials is −3.59, indicating that the two inputs are strong complements. The materials variable in the dataset is primarily composed of fertilisers and pesticides. While Hungarian agriculture has witnessed a secular increase in all input prices over time, some of the largest increases registered have been in energy, fertiliser and pesticide prices. Table 4 provides an idea of the scale of price increases for these inputs,[6] while Table 5 indicates how the input quantities have changed over time. The huge reductions in the use of fuels, fertilisers and pesticides as observed in Table 5 have doubtless been caused to a significant extent by the changes in the *own prices* of these inputs as reported in Table 4. Another routinely discussed reason is the removal of *output price* subsidies. The analysis here indicates a third reason, one that has not been discussed before, i.e. the *cross-price* effects indicating complementarity of the inputs. It appears that the evolution of input prices has in some ways been fortuitous for Hungarian agriculture. The prices of energy and materials have increased tremendously over the last decade, more due to the removal of subsidies in the transition process rather than as an explicit effort to conserve energy or curtail polluting input (fertilisers and pesticides) use. If these categories of inputs had been substitutes, the parallel increases in the prices of these input categories would have resulted in cross-price effects working against each other. However, complementarity between energy and materials implies that the cross-price effects have worked in tandem to restrict the use of energy and materials. Energy price increases discourage energy and materials use, and similarly materials price increases also discourage energy and materials use.

The AES estimates for energy and capital (7.98) and energy and labour (4.58) are both large and positive, indicating that the pairs are strong substitutes. Unfortunately, price and volume series for labour and capital over the 1990s do not seem to be readily available. Volume and price series for investments are available in the Hungarian Statistical Yearbook of Agriculture. These however, include hunting and forestry besides agriculture. Additionally, the price index calculation methods change halfway through the sample because of the introduction of VAT. No wage index is available, and the only information on labour quantity is the number of people employed in agriculture. Discussions in various reports on Hungarian agriculture however, indicate that the trends for these inputs have been similar to those for energy and materials, that is, increasing prices and declining input use, although perhaps the price changes have not been as drastic. The elasticity of substitution between energy and capital has been the subject of a major debate in the literature over the years, comprising at least 50 studies (Thompson and

[6] As far as data on input use and prices over the 1990s for the agricultural sector as a whole are concerned, 'fuels and lubricants' is the closest available proxy for 'energy'. In fact, it proved surprisingly difficult to get useful data on input use for the agricultural sector as a whole, although this is improving as new accounting systems at the national and firm level are put in place.

Taylor, 1995). After careful analysis of the evidence, Thompson and Taylor (1995) conclude that the evidence points towards capital and energy being substitutes for each other. This conclusion is supported by these estimates for Hungarian agriculture. Once again, this relationship is somewhat fortuitous for Hungary, for if the two inputs were complements, the surging energy prices over the last decade would have curtailed capital formation and thereby the long-run growth of the sector. The large AES estimate of 7.98 for energy and capital indicates that the energy price increases have actually worked to encourage capital formation in Hungarian agriculture. Although this conclusion may appear odd in view of the capital base shrinkage on Hungarian farms over the last decade, it must be remembered that no elasticity of substitution is a sufficient statistic for predicting input use, particularly when several prices, input as well as output, are changing at the same time. Perhaps the ideal way to phrase the result would be to say that the strong substitutability has prevented the capital formation situation from being even worse. A similar interpretation is appropriate for the AES of 4.58 between energy and labour. Here, the high level of substitutability may have prevented a worse agricultural employment situation than has actually taken place.

The MES, being a 'two-factor one-price' elasticity, yields further insights, particularly when used in combination with AES estimates (Chambers, 1991). The MES between factors i and j can be interpreted as the change in the ratio i/j in response to a 1% change in the price of j. Therefore the MES is asymmetric in that the MES between i and j is different from the MES between j and i. AES complements may be classified as MES substitutes. To illustrate the logic behind this, consider the capital/labour elasticities. The capital/labour AES is -2.1, indicating Allen complementarity, and so an increase in the wage rate results in a decline in capital formation. The MES of 0.59, however, implies that capital and labour are Morishima substitutes. The logic is as follows: an increase in the wage rate results in a decline in labour use (implied by the concavity of the production function). However, it also results in a decline in capital use, since labour and capital are Allen complements. Thus both the numerator and the denominator in the capital/labour ratio are declining. In this case, the own price effect (labour reduction) outweighs the cross-price effect (capital reduction), resulting in the capital/labour ratio increasing.

However, in the case of the energy–materials pair the AES classification as complements holds in the MES case as well, since both MESs for the pair are negative. This result highlights a point made earlier, that energy use in Hungarian agriculture is extraordinarily sensitive to the prices of other inputs. The AES indicates that increased materials prices would discourage energy use. It would, of course, also discourage materials use. However, the estimates indicate that the cross-price effect is even stronger than the own price effect.

Finally, the MES between capital and energy helps shed light on an old energy policy question. Is energy conservation better promoted by energy price increases, or by tax credits (subsidies) for energy-saving capital? The capital/energy MES in response to an energy price change (1.56) is considerably larger than the energy/capital MES in response to a capital price change (0.56). This implies that continued energy price increases would better accomplish the task of reducing energy consumption and promoting investment in energy-saving machinery than would the capital subsidy approach.

8. Conclusion

This paper uses a unique panel dataset of Hungarian agriculture in the pre/early-transition period to measure the direction and extent of energy input substitution. The results are used to derive implications for energy-related issues in the Hungarian agricultural sector in the late/post-transition period. Three specific issues are analysed: the relationship between relative input price changes and energy use, the effect of rising energy prices on capital formation, employment and chemical input use, and the question of how policy may be best designed to make the Hungarian agricultural sector less energy intensive.

The methodology is based upon a direct estimation of technology, thereby avoiding behavioural assumptions that may be inappropriate for the Hungarian economy in early transition. Recently developed econometric methods based on panel techniques, GMMs, and a series of instrument exogeneity tests are used, and the strong exogeneity model is chosen on this basis. Then, estimates of both Allen and Morishima elasticities are constructed.

A number of results follow. Firstly, the AES estimates involving energy are the largest among the set of all inputs pairs. This suggests that energy use in Hungarian agriculture is most responsive to the price of other inputs and points out an alternative avenue for effective energy policy, indirect price control. Energy and chemical inputs (materials) are found to be strong complements, implying that the sharp and sustained increases in the prices of these inputs over the 1990s have worked jointly to restrict energy and chemical input use. With respect to capital and energy, the results confirm the conventional wisdom that these inputs are substitutes, and therefore that energy price inflation has and will enhance rather than inhibit capital formation, output held constant. A similar result holds good for the effect of energy price inflation on agricultural employment.

The MES estimates allow further inferences. The estimated pair of MES between energy and materials serve to highlight the strength of the cross-price effects between the pair, which overshadow the own price effects. The pair of MES estimates between energy and capital indicate that Hungary's policy of energy-use curtailment and promotion of energy-saving capital may be better promoted by energy price increases than by capital subsidies.

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