ACTA AGRONOMICA ÓVÁRIENSIS VOL. 53. NO. 1.

Supply response on the Hungarian pork sector

ZOLTÁN BAKUCS[1] – RICHÁRD MÁRKUS[2]

[1] Hungarian Academy of Science
Institute of Economics
Budapest

[2] University of West Hungary
Faculty of Agricultural and Food Sciences
Institute of Business Economics and Management Sciences
Mosonmagyaróvár

SUMMARY

Despite of the increasing production and consumption of white meats, swine production is still one of the most important sectors of animal husbandry worldwide as well as in the European Union. In Hungary over recent decades, the swine industry has undergone significant changes. The livestock has decreased sharply from more than 8.5 million in 1989 to 3.3 million currently. After the post 1989 increase of herd size on family farms, their share diminished, and currently two-thirds of output is produced by corporate farms. It appears that small scale pork producers have major difficulties, and must consider all cost reducing alternatives to improve their competitiveness. Considering the pressure on purchase prices from the downstream markets, and the fact that feed represents about 50–60% of total production costs, in this paper we analyze the influence of these two factors on swine producer's supply response. We employ Vector Error Correction Model specification, following the theoretical model of *Hallam* and *Zanoli* (1993). Estimated long-run elasticities highlight farmers' reliance on the slaughter purchase price of live pigs and soybean meal price.
Keywords: error correction model, supply response, pork sector, Hungary.

INTRODUCTION

Swine production was one of the most dramatically affected segments of Hungarian agriculture by the massive changes after 1989. The dismantling of socialist agricultural enterprises and cooperatives, and the shift towards private farming, had a major impact upon the pig stock which was reduced by half in just a couple of years. Newly emerged family farms producing pork are fighting for survival, many being subsistence farms.

There are several reasons for the decreasing Hungarian pig production including the loss of former markets, enhanced competition after joining to EU, inefficient pig production (daily weight gain, feed conversion rate), high price of feedstuff, etc.

The Hungarian meat industry is characterized by a distorted market structure, emphasized by the large number of small, inefficient firms. The dramatic decrease of raw material production left many of the formerly efficient larger size companies struggling with unused processing capacity. *Jansik* (2000) studying the foreign direct investment (FDI) in Hungary, finds that industries characterized by a monopolistic market structure (sugar, vegetable oil, tobacco, soft drinks, starch) were privatized in the early 1990s, having over 70% foreign ownership of their capital. Meat processing is the largest segment of the food industry, accounting for over 18% of the total Hungarian food processing output. The sales of meat industry indicate a slightly growing trend. The number of firms decreased by about 50% between 1996 and 2000, and then it started to grow again. The privatization of the meat industry started late, in the mid 1990s, and was characterized by low FDI. In 2005, approximately 40% of total capital was in foreign ownership. Thus, the concentration process was delayed; the five firm concentration ratios in the meat industry are still rather low with 30.6% in 1992, and 44.1% in 2003. The Hungarian pork sector has experienced numerous structural changes in the past 15 years. From 9.5 million head in September 1990, the pig stock decreased to 4.3 million by December 1994, and it has fluctuated at around 5 million head ever since. One important feature of the Hungarian swine industry is the large number of small-scale farms. Even before privatization, small-scale farms accounted for 50% of the total pig stock, a figure that has not changed significantly since 2005. Many of these small-scale farms do not have commercial activity, i.e. they are subsistence farms. However, a large proportion of them sell their products, forming a two-tier commercial and family pork production system. The average herd size by farm type illustrates unambiguously a dual production structure in the Hungarian pork sector. The average herd size varies between 9–16 pigs. However these numbers hide the significant differences between various types of farms. Private farms on average hold 5 to 7 pigs, while the average herd size for economic organizations is 3.3 to 4.4 thousand pigs. Considering the technology of pig fattening, there is a significant fluctuation concerning the most important feed ingredients (soybean, corn, wheat, barley). At present, soybean meal is the most important feed ingredient. Despite the decreasing animal livestock, Hungary is still a net importer of soybean meal with approximately 600–700 thousand tonnes being imported every year.

The remainder of this manuscript is organized where the second section briefly reviews the existing supply response literature in Central and Eastern European (CEE) countries, followed by the empirical methodology in section 3. Variables are described in section 4, while results are presented in section 5. The last section summarizes and offers some conclusions on the implications for the Hungarian pork producers.

Literature review

There is already a great wealth of literature examining various aspects of the transition period from the transformation in the farm structure to competitiveness and efficiency analysis or vertical price transmission of various sectors. Research into the key determinants, and indeed the estimation of an appropriate agricultural supply response model for transition economies is, however, scarce. One reason for the limited availability of data is that only semi-annual pig stock data exists. Of the papers focusing on supply response in CEE countries, *Hallam* (1998) analyses the supply response in some transition economies, namely Bulgaria, Romania and Slovenia. The author points out the problems of estimating econometric supply models due to the numerous structural breaks that occurred during the transition period, and the lack of sufficiently long time series data. *Mishev et al.* (1998) estimate the price elasticities of supply for Bulgarian crop products, concluding higher own price elasticities than in developed economies, mostly determined by input shortages. *Nyárs* and *Vizvári* (2005) apply linear and non-linear regression equations to estimate the supply response on the Hungarian pork market. The authors estimate that in good market conditions for pork producers (low input, high output prices), the Hungarian pork sector can produce 526,000 tons of live pigs for slaughter, while in unfavourable market environments (high input, low output prices) the capacity is reduced to 411,000 tons.

Contrary to the *Nyárs* and *Vizvári* (2005), the aim of this paper is to estimate a parsimonious Vector Error Correction econometric model, using yearly data from the past 21 years. We follow the methodology outlined in *Hallam* and *Zanoli* (1993) which prove the superiority of error correction specifications to the more common partial adjustment models with regard to agricultural supply response. Earlier studies (e.g. *Ness* and *Colman* 1976, *Holt* and *Johnson* 1988 or *Hallam* and *Zanoli* 1993) demonstrated that the target breeding herd may be modelled as a linear function of own price (pig purchase price) and feed price. Also, the pig production technology excludes the possibility of significant cross-price elasticities with respect to other outputs.

Materials and methods

The long-run supply function may be modelled as:

$$st = c + \beta 1pe + \beta 2fe \qquad (1)$$

where st is the target breeding stock, pe is the expected real purchase price, and fe is the expected feed price. There are several possibilities of defining farmers' price expectations, naive, rational or adaptive. *Hallam* and *Zanoli* 1993 shows, that a VECM model can adequately describe real pig and feed prices through autoregressive lags. Thus, we wish to estimate the long- and short-run pig supply elasticities with respect to slaughter pig and feed price.

The empirical procedure is based on modern time series econometrics, namely Vector Error Correction model estimations. Series are first tested for unit roots, then cointegration, followed by the estimation of a Vector Error Correction Model simultaneously depicting both long and short run response of the breeding stock to changes in pork purchase and soybean meal prices.

Testing for Unit Roots

With time series data, one needs to pay a particular attention to the stationarity of the variables. In the presence of unit roots, classical ordinary least square (OLS) regression yields biased estimates, invalid tests, and ultimately, spurious regressions. Considering the first order autoregressive process, AR(1):

$yt = \rho yt-1+et$ where $t = \ldots, -1, 0, 1, 2, \ldots$, and et is white noise error stochastic term. (2)

The process is considered as stationary, if $|\rho| < 1$, thus testing for stationarity is equivalent with testing for unit roots ($\rho = 1$). (2) is rewritten to obtain

$$\Delta yt = \delta yt-1+et \quad \text{where } \delta = 1-\rho \tag{3}$$

and thus the test becomes:

null hypothesis $H0 : \delta = 0$ against the alternative hypothesis $H1: \delta < 0$.

There are a large number of unit root testing procedures in the literature, see *Maddala* and *Kim* (1988) for a detailed discussion. Considering the notoriously low size and power properties of unit root tests, in this paper we employ three unit root tests that have alternative null hypotheses. The null hypothesis of the Augmented Dickey-Fuller, ADF (*Dickey* and *Fuller* 1979) and Phillips-Perron, PP (*Phillips* and *Perron* 1988) test is a unit root in the variable against the alternative of stationarity. The KPSS (*Kwiatkowski et al.* 1992) procedure tests the null of stationarity against the alternative of a unit root process.

Cointegration analysis and Vector Error Correction Modelling

Non stationary variables may be analyzed in a cointegration framework. We test for cointegration using Johansen's multivariate cointegration approach (*Johansen* 1988). This procedure is a Maximum Likelihood (ML) approach in a multivariate autoregressive framework with enough lags introduced to have a well-behaved disturbance term. It is based on the estimation of a Vector Error Correction model (VEC) of the form:

$$\Delta Zt = \Gamma 1\Delta Zt-1+ \ldots+\Gamma k-1\Delta Zt-k+1+\Pi Zt-k+ut \tag{4}$$

where $Zt = [st, pt, sft]'$ is a (3x1) vector containing the three I(1) variables, where s stands for Hungarian breeding stock (sow stock), p the log of pork producer purchase price and sf the price of feed, t for time period, $\Gamma 1, \ldots \Gamma k+1$ are vectors of the short-run parameters, Π is matrix of the long-run parameters, and ut is the white noise stochastic term. Monthly seasonal dummy variables may also be included.
$\Pi = \alpha\beta`$, where matrix α represents the speed of adjustment to disequilibrium and β is a matrix which represents up to (n–1) co integrating relationships between the non-stationary

variables. There are five possible models in (4) depending on the deterministic specification. Following *Harris* and *Sollis* (2003) these 1 to 5 models are defined as: (M1) no intercept or trend is included; (M2) the intercept is restricted to the cointegration space; (M3) unrestricted intercept without trends; the intercept in the cointegration space is combined with the intercept in the short-run model resulting in an overall intercept contained in the short-run model; (M4) if there exists an exogenous linear growth not accounted for by the model, the cointegration space includes time as a trend stationary variable; and (M5) allows for quadratic trends in Zt.

Dataset and Descriptive statistics of variables

Annual data between 1981 and 2009 were provided by the *Hungarian Central Statistical Agency (HCSA).* The dataset consists of the sow stock, s (used as proxy for the breeding stock), annual average purchase price of live pigs for slaughter, p, and the price one of the most important feed ingredients used in Hungarian pig production systems, soybean meal, sf. Price data was deflated to 1989 by the national Consumer Price Index. *Figures 1., 2. and 3.* present the log of the sow stock, pig purchase price and soybean meal price, respectively. The first graph illustrates the dramatic fall in the Hungarian total stock and indeed breeding stock after the fall of the socialist regime, discussed in the market overview section of this paper. In real terms, however, price data also appears to be downward trended.

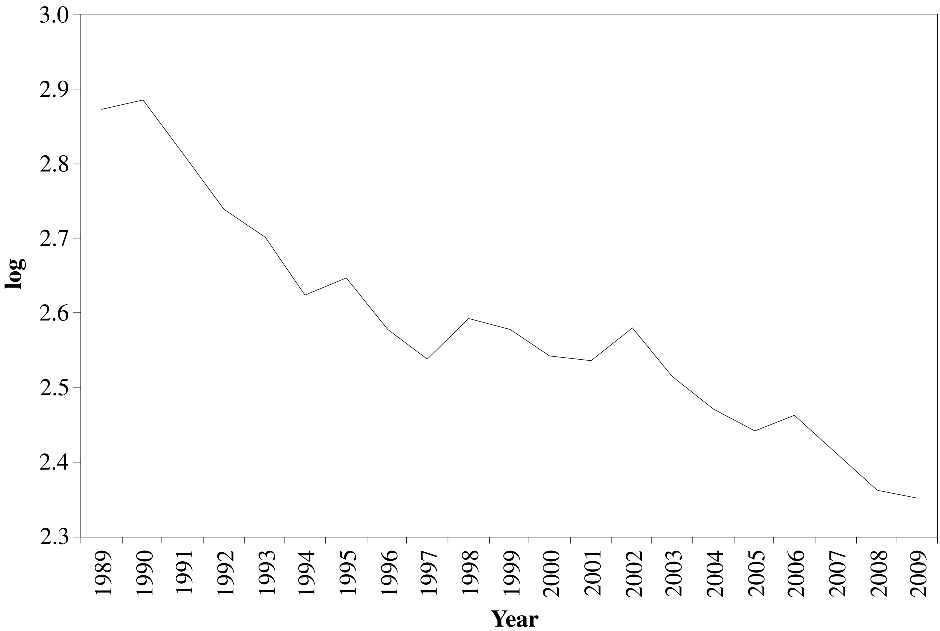

Figure 1. The log of Hungarian sow stock

Note: Own calculations, data provided by *HSCA*

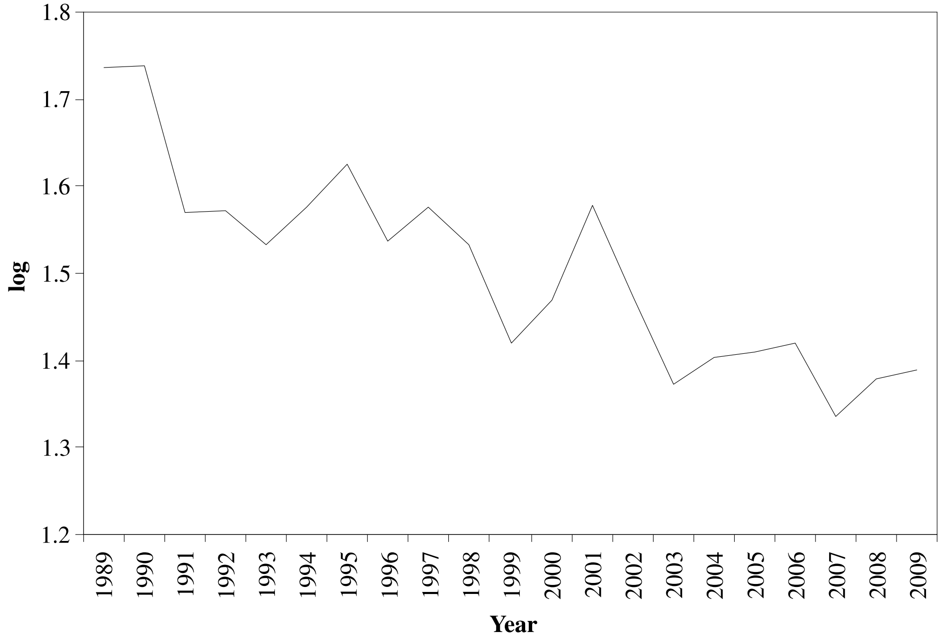

Figure 2. The log of pork producer purchase price
Note: Own calculations, data provided by *HSCA*

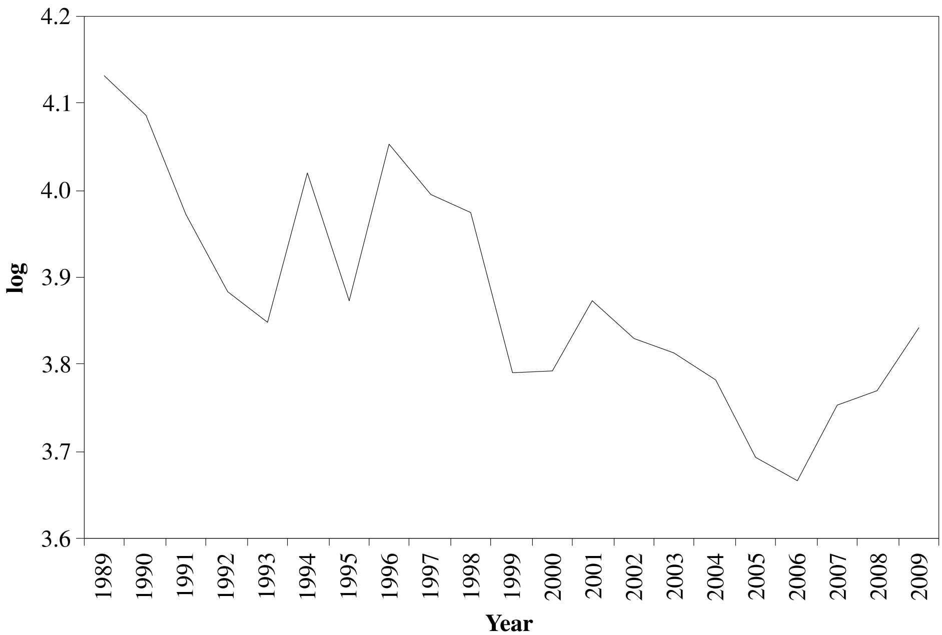

Figure 3. The log of soybean meal price
Note: Own calculations, data provided by *HSCA*

Table 1. presents the descriptive statistics of the 3 time series.

Table 1. Descriptive statistics of variables

Variable	No. of Observations	Mean	Std. Dev.	Min	Max
s	21	2.583	0.148	2.352	2.885
p	21	1.506	0.115	1.335	1.737
sf	21	3.878	0.125	3.667	4.131

Note: Own calculations, data provided by *HSCA*

Evaluation of results

To determine the properties of the time series data, a battery of unit root tests were applied. The null hypothesis of ADF and PP tests is unit root, against the alternative hypothesis of stationary series. The null hypothesis of the KPSS test however is stationary series against the alternative hypothesis of unit root in the series. *Table 2.* presents unit root test results in the sow stock, pig purchase price and soybean meal price series. The upper panel of *Table 2.* presents test statistics (with 5% critical values below each statistic in brackets) where the test regression contains an intercept only. Results obtained by test regressions with intercept and trend as deterministic specifications are presented in a similar fashion in the lower panel of *Table 2.*

Table 2. Unit root tests

Test statistic (5% crit. value)	s	p	sf
with intercept only			
ADF	–0.462	1.673	–2.454
	(–3.144)	(–3.119)	(–3.02)
PP	–1.804	–2.096	–2.433
	(–3.02)	(–3.02)	(–3.02)
KPSS‡	0.616**	0.63**	0.546**
	(0.463)	(0.463)	(0.463)
with intercept and trend			
ADF	–4.251**	–1.356	–1.084
	(–3.791)	(–3.875)	(–3.875)
PP	–2.06	–6.899***	–2.946
	(–3.658)	(–3.658)	(–3.658)
KPSS‡	0.132	0.322***	0.074
	(0.146)	(0.146)	(0.146)

Note: *** significant at 1%, ** significant at 5%, * significant at 10%
‡ the null hypothesis of the KPSS test is that variable is stationary

Mixed results were obtained. With intercept only, PP and ADF tests cannot reject the unit root null hypothesis, whilst the KPSS test significantly rejects the stationarity null hypothesis.

Figures 1., 2. and 3. however suggest that data might be trended but with intercept and trend specification, the picture is less clear. For the sow stock variable, ADF and KPSS tests suggest that data is stationary, while the PP test does not reject the unit root null. For the purchase price series, PP test rejects the unit root null, but ADF and KPSS tests suggest that the data contain unit root. Similarly, for the soybean meal price, KPSS suggest that data is stationary, however ADF and PP tests conclude that data is not stationary. Considering the notoriously low power and size properties of unit root tests, we carefully conclude that all three time series are not stationary, i.e. integrated of order one, I (1).
Non-stationary data must be cointegrated in order to estimate any long-run relationship between variables. A number of different deterministic specifications were sequentially tested for co integration. Test results for models M2 and M3 (see the methodology section) are presented in *Table 3.*

Table 3. Johansen cointegration tests

Number of CI vectors	**P-value** (intercept only)	**P-value** (intercept and trend)
0	0.029	0.007
1	0.478	0.086
2	0.382	0.376

Note: 3 lags in first differences were selected by AIC criteria

The null hypothesis of no cointegration between the 3 variables is rejected for both specifications. With trend included in the cointegration space, even the one cointegrating vector null hypothesis may be rejected in favour of 2 vectors at 10% level of significance. Based on the results from *Table 3.*, we consider the sow stock, pork purchase price and soybean meal variables co integrated with 1 cointegration vector. The long-run relationship between these variables, basically the supply response function is:

$$s = 2.758p - 1.391sf + 3.794 \quad (5)$$

Since data is in logs, coefficients represent long-run elasticities of the sow stock (breeding herd) with respect to the pork purchase price and soybean meal price. Thus 1% increase in expected pig purchase price induces an increase of 2.75% of the breeding stock, while the 1% increase in the expected feed price decreases the breeding stock by 1.39%. The estimated VECM model of the pork supply response is presented in *Table 4.* with some diagnostics and coefficients of determination in the lower panel. The upper panel contains the long-run supply response (with t statistics in brackets) identical to *Equation 4.* The middle panel presents the short-run dynamics of the VECM also meant to model the rational expectations hypothesis of the pig farmer through the autoregressive lags of variables. The first row of the middle panel contains the coefficients of the error correction term, (α in *Equation 3*, see methodology section) and their corresponding t statistics. These coefficients measure the speed of adjustment towards the long-run equilibrium, i.e. how fast the system returns to its long-run equilibrium path should an exogenous shock occur.

Table 4. Supply response Vector Error Correction Model

Variable	Cointegration equation		
st-1	1.000		
pt-1	–2.758		
	(–14.259)		
sft-1	1.391		
	(8.906)		
C	–3.794		
	(–10.882)		
short-run dynamics			
	Δst	Δpt	Δsft
error correction	0.271	0.478	–1.065
	(1.809)	(2.090)	(–2.364)
Δst-1	-0.162	–1.014	–0.004
	(–0.707)	(2.887)	(–0.006)
Δst-2	0.475	0.920	0.296
	(2.838)	(3.597)	(0.587)
Δpt-1	1.017	1.224	–1.438
	(3.215)	(2.531)	(1.508)
Δpt-2	0.257	0.461	–0.712
	(1.239)	(1.456)	(1.140)
Δsft-1	–0.142	–0.418	0.271
	(–1.001)	(1.923)	(0.632)
Δsft-2	0.052	–0.200	0.106
	(0.717)	(1.792)	(0.483)
diagnostics‡			
LM test AR(1)	0.106		
LM test AR(2)	0.124		
Jarque-Bera	0.706	0.921	0.148
R2	0.601	0.659	0.386

‡ p values (significance) are presented

The coefficients of adjustment are highly significant in the own price and feed price equations, but surprisingly only marginally significant (at 10%) for the breeding stock equation. A non-significant coefficient would mean that the short-run equation does not adjust to deviations from the long-run equilibrium, i.e. it is weakly exogenous on long-run. The model appears to be well specified, the null hypothesis of no first and second order autocorrelation in the residuals cannot be rejected at 5% level of significance. The residuals are normally distributed, while the coefficients of determination range between 38 and 66% acceptable for this kind of analysis.

Conclusions

The Hungarian swine industry, production as well as processing, has undergone extraordinary changes during the past two decades. The increasing competition and the continuous changes in the structure of farming and industry revealed several problems of the pork production in Hungary. In this paper, we showed that there is a long-run cointegrating relationship between the size of breeding stock, pork purchase price and soybean meal price. The analysis revealed the relatively high importance of expected pig purchase price and the price of the most important input, the price of soybean meal in the production decisions of farmers. Estimated long-run elasticities of the sow stock with regards to the pork purchase price and soybean meal are quite high; a 1% increase in expected pig purchase prices induces a 2.75% increase of the breeding stock, while a 1% increase in the expected soybean meal price decreases the breeding stock by 1.39%.

Kínálati reakció alakulása a magyarországi sertéshús szektorban

ZOLTÁN BAKUCS[1] – RICHÁRD MÁRKUS[2]

[1] Magyar Tudományos Akadémia
Közgazdaság Tudományi Intézet
Budapest

[2] Nyugat-magyarországi Egyetem
Mezőgazdaság- és Élelmiszertudományi Kar
Vállalatgazdasági és Vezetéstudományi Intézet
Mosonmagyaróvár

Összefoglalás

A sertéstenyésztés még mindig a legfontosabb állattenyésztési ágazat annak ellenére, hogy a baromfihúsok termelése és fogyasztása folyamatosan növekszik mind világviszonylatban, mind az Európai Unióban. A sertéságazaton belül jelentős változás volt megfigyelhető hazánkban az elmúlt évtizedek során. Az állomány drasztikus csökkenése szemlélteti, hogy az 1989-ben még több mint 8,5 milliós állomány napjainkra mindösszesen 3,3 millióra tehető. A rendszerváltás utáni években a hazai sertésállomány jelentős részét képezte az úgymond „háztáji" sertésállomány, napjainkban azonban az állomány kétharmadát gazdasági társaságokon belül állítják elő. Ez a tendencia is szemlélteti, hogy a kisgazdaságoknak minden lehetséges költségcsökkentő alternatívát figyelembe kell venniük versenyképességük javítása érdekében. A sertéstenyésztők döntési reakcióit – ezáltal a hazai sertésállomány változását – vizsgáltuk az értékesítési piac által diktált gyakran nyomott árak, továbbá a takarmányozási költségek – az összes költségen belül elérik az 50–60 százalékot – függvényében. *Hallam* és *Zanoli* (1993) alapján a vektorhiba korrekciós modellt alkalmaztuk.

A becsült hosszú távú elaszticitás alapján előrevetíthetjük a termelők gazdasági döntéseit a vágósertés felvásárlási árak és a szója árak változásainak hatására.
Kulcsszavak: hiba korrekciós modell, kínálati reakció, sertéshús szektor, Magyarország.

References

Dickey, D. A. – Fuller, W. A. (1979): Distributions of the Estimators For Autoregressive Time Series With a Unit Root. Journal of the American Statistical Association **75,** 427–431.

Hallam, D. – Zanoli, R. (1993): Error correction models and agricultural supply response. European Review of Agricultural Economics **20,** 151–166.

Hallam, D. (1998): Analysing Agricultural Supply Response in Economies in Transition. Options Méditerranéennes. Série B. Etudes et Recherches **22,** 123–131.

Harris, R. – Sollis, R. (2003): Applied Time Series Modelling and Forecasting. John Wiley & Sons Ltd., Chichester, England.

Holt, M. T. – Johnson, S. R. (1988): Supply dynamics int he US hog industry. Canadian Journal of Agricultural Economics **36,** (3) 313–335.

Jansik, Cs. (2000): Determinants and Influence of Foreign Direct Investments in the Hungarian Food Industry in a Central and Eastern European Context. Agrifood Research Finland Economic Research (MTTL) Publications **102.**

Johansen, S. (1988): Statistical Analysis of Cointegrating Vectors. Journal of Economic Dynamics and Control **12,** 231–254.

Kwiatkowski, D. – Phillips, P. C. B. – Schmidt, P. – Shin, Y. (1992): Testing the null hypothesis of stationary against the alternative of a unit root. Journal of Econometrics **54,** 159–178.

Maddala, G. S. – Kim, I. (1988): Unit Roots, Cointegration, and Structural Change. Cambridge University Press, Cambridge.

Mishev, P. – Tzoneva, M. – Ivanova, N. (1998): Supply Response of Bulgarian Agriculture Over the Transition Period. Options Méditerranéennes. Série B. Etudes et Recherches **22,** 147–154.

Ness, M – Coleman, D. R. (1976): Forecasting the size of UK pig breeding herd. Bulletin **157,** University of Manchester, Department of Agricultural Economics.

Nyárs L. – Vizvári B. (2005): On the Estimation of the Supply Function of the Hungarian Pork Market. Journal of Central European Agriculture **6,** (4) 521–530.

Phillips, P. C. B. – Perron, P. (1988): Testing for a unit root in time series regression. Biometrika **75,** 335–346.

Address of the authors – A szerzők levélcíme:

MÁRKUS Richárd
University of West Hungary
Faculty of Agricultural and Food Sciences
H-9200 Mosonmagyaróvár, Vár 2.
E-mail: richardmarkus.nyme@gmail.com