The Influence of Macroeconomic Variables on the Hungarian Agriculture

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Abstract

This paper focus on the time adjustment paths of the exchange rate and prices in response to unanticipated monetary shocks following model developed by Saghaian et al. (2002). We employ Johansen's cointegration test along with a vector error correction model to investigate whether agricultural prices overshoot in a transition economy. The empirical results indicate that agricultural prices adjust faster than industrial prices to innovations in the money supply, affecting relative prices in the short run, but strict long-run money neutrality does not hold.
Keywords: agricultural prices, exchange rates, monetary shocks, overshooting, transition economy

JEL classification: C32, E51, P22, Q11

1. Introduction

There is a continuously growing literature on the agricultural transformation in Central an Eastern European countries (see survey Brooks and Nash 2002; Rozelle and Swinnen 2004). The research has focused on various aspects of transition, including land reform, farm restructuring, price and trade liberalisation and etc. However, until now macroeconomic aspects of agricultural transition were neglected. The agricultural economics literature has emphasised the importance of macroeconomics and financial factors in the determination of agricultural prices already in the second half of eighties (e.g. Bessler, 1984; Chambers, 1984; Orden, 1986a,b; Devadoss and Meyers, 1987; Orden and Fackler, 1989). Recently there has been renewed interest in the analysis of impact of monetary variables for agricultural prices (Zanias 1998; Saghaian et al, 2002; Ivanova et al. 2003; Cho et al., 2004; Peng et al., 2004) employing cointegration and Vector Error Correction (VEC) framework. Previous empirical research based on mainly U.S. agriculture suggests that any changes in macroeconomic variables should have an impact on agricultural prices, farm incomes and agricultural exports. Therefore, it is reasonable assume that a transition country characterised less stable macroeconomic environment these effects are more profound. Surprisingly, the interest has been almost non-existent in Central-Eastern Europe, except Ivanova et al. (2003), who studied the macroeconomic impacts on the Bulgarian agriculture.

Monetary policy has real and nominal effects on the overall economy and the agriculture in short run and medium run, but generally no real effects in long run (Ardeni and Freebairn, 2002). There are number of direct linkages between monetary policy and agricultural sector. However, in this study we focus exclusively on the overshooting hypothesis claiming that monetary changes can have real short-run effects on the prices of agricultural commodities. This indicates that money supply is not neutral and monetary impacts can change relative prices in the short run. The paper examines the short-run overshooting of agricultural prices in Hungary using cointegration and VEC framework. The empirical results have also implications for long-run money neutrality. This issue is important in transition countries, because price variability is much less for industrial prices then for agricultural prices during the transition period especially comparing similar price movements in developed countries. Overshooting of agricultural prices can at least partially explain the observed agricultural-price variability. These monetary impacts and financial factors have policy implications as well. The short- and long-run impacts of monetary policy have been very important for the Hungarian agricultural sector due to lack of credibility of farm policy, where farm incomes are much more influenced by market prices. If money is neutral in the long run, commodity price overshooting can still have significant effects on short-run farm income and the financial viability of farms.

The paper is organised as follows. Section 2 discusses theoretical background and related empirical evidence. The time series methodology employed is described in section 3. The data and the results of empirical models are presented in the section 4. Finally, the conclusions and implications of the results on the Hungarian agriculture are drawn in the last section.

2. Theoretical considerations and empirical evidence

At least since Schuh (1974) interest has continued in the possible impacts of monetary policy on agricultural markets. This issue is important because policies to stabilise agricultural markets should consider the sources of volatility within agri-food sector. The main issue is that whether levels of agricultural and non-agricultural prices respond proportionally to changes in the level of money supply in the long run, and whether money is neutral in short run. Various explanations are available for relative price movements. It is usually assumed that agriculture is a competitive sector in which its prices are more flexible than in non-agricultural (fix price) sectors. Consequently, expansionary monetary policy favours agriculture, because farm prices can be expected to increase faster than non agricultural prices, while restrictive monetary policy shifts prices against agriculture. Bordo (1980) argues that agricultural commodities tend to be more highly standardised and therefore exhibit lower transaction costs than manufactured goods. Consequently, agriculture is characterised rather short term contracts which lead a faster response to a monetary shock. Alternatively, Tweeten

(1980) argue that price shocks stemming in oligopolistic non-agricultural sector and accommodated by expansionary monetary policy, cause inflation and place agriculture in a price-cost squeeze.

Other streams of research address the broader macroeconomic environment. Arising from Dornbusch's (1976) overshooting models of exchange rate determination, these studies establish the linkages among exchange rates, money, interest rate and commodity prices. Frankel (1986) applied Dornbusch's model in which exchanges rates, money supply, interest rate and aggregate demand determine commodity prices assuming closed economy. He emphasised the distinction between "fix-price" sectors (manufacturers and services sector), where prices adjust slowly and "flex-price" sector (agriculture), where prices adjust instantaneously in response to a change in the money supply. In Frankel's model, a decrease in nominal money supply is a decline in real money supply. This leads to an increase in interest rate, which in turn depresses real commodity prices. The latter then overshoot (downward) their new equilibrium value in order to generate expectation of a future appreciation sufficient to offset higher interest rate. In the long run, all real effects vanish. Lai, Hu and Wang (1996) employed Frankel's framework and phase diagram to investigate how money shocks influence commodity prices. They found that with unanticipated monetary shocks, commodity prices overshoot, but, if manufactured prices respond instantly, commodity prices undershoot. Saghaian, Reed and Marchant (2002) extended Dornbusch's model with agricultural sector and allowing for international trade of agricultural commodities. Agricultural prices and exchange rate are assumed flexible, while industrial prices are assumed to be sticky. Employing small open country assumption, they showed that when monetary shocks occur, the prices in flexible sectors (agriculture and services) overshoot their long-run equilibrium values. Furthermore, they showed that with presence of a sticky sector, in case of monetary shock, the burden of adjustment in the short run is shared by two flexible sectors and having a flexible exchange regime decreases the overshooting of agricultural prices and vice versa. The extent of overshooting in the two flexible sectors depends on the relative weight of fix-price sector.

All studies found significant effects of changes in macroeconomic variables for monetary policy and exchanges rates in the short run. Several authors found that farm prices respond faster than non farm prices, which consistent with hypothesis that relative prices change as money supply changes due to price level in the various sectors change differently (Bordo 1980, Chambers 1984, Orden 1986a and 1986b, Devadoss and Meyers 1987, Taylor and

Spriggs, 1991, Zanias 1998, Saghaian, Reed and Marchant 2002). However, Bessler (1984), Grennes and Lapp (1986) Robertson and Orden (1990), and Cho et al. (2004) found that relative agricultural prices are not affected by nominal macroeconomic variables. These studies also show that although short run effects of money changes may be different, long run effect are equal supporting the long-run neutrality of money (Ardeni and Rausser 1995). However, Saghaian et al. (2002) results reject the hypothesis of the long-run neutrality of money. It should be noted that these results should be interpreted only with care. First, time-series studies of links between the agriculture and the rest of economy are often sensitive to variable choices. Second, as Ardeni and Freebairn (2002) pointed out, many studies lack an appropriate treatment of the time series properties of data implying misleading results especially on the case of earlier research. Finally, the main feature of the literature is that many studies do not relate directly a specific macroeconomic model, except Saghaian et al. (2002), rather they use a set of explanatory variables suggested by previous studies.

3. Empirical Procedure

Even as many individual time series contain stochastic trends (i.e. they are not stationary at levels), many of them tend to move together on long run, suggesting the existence of a long run equilibrium relationship. Two or more non-stationary variables are cointegrated if there exists one or more linear combinations of the variables that are stationary. That implies that the stochastic trends of the variables are linked over time, moving towards the same long-term equilibrium.

3.1. Testing for unit roots

Consider the first order autoregressive process, AR(1):

$y_t = \rho y_{t-1} + e_t \quad t = \ldots,-1,0,1,2,\ldots$, where e_t is White Noise. (1)

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

(1) is rewritten to obtain

$\Delta y_t = \delta y_{t-1} + e_t$, where $\delta = 1 - \rho$ (2)

and thus the test becomes:

$H_0 : \delta = 0$ against the alternative $H_1 : \delta < 0$.

Maddala and Kim (1998) argues, that because of the size distortions and poor power problems associated with the Augmented Dickey-Fuller unit root tests, it is preferable to use the DF-GLS unit root test, derived by Elliott, Rothenberg and Stock (1996).

Elliott, Rothenberg and Stock develop the asymptotic power envelope for point optimal autoregressive unit root tests, and propose several tests whose power functions are tangent to the power envelope and never too far below (Maddala and Kim, 1998). The proposed DF-GLS test works by testing the $a_0=0$ null hypothesis in regression (3):

$$\Delta y^d_t = a_0 y^d_{t-1} + a_1 \Delta y^d_{t-1} + \ldots + a_p \Delta^d_{t-p} + e_t \tag{3}$$

where y^d_t is the locally detrended y_t series that depends on whether a model with a drift or linear trend is considered. In case of a model with a linear trend, the following formula is used to obtain the detrended series y^d_t:

$$y^d_t = y_t - \hat{\beta}_0 - \hat{\beta}_t\ . \tag{4}$$

$\hat{\beta}_0$ and $\hat{\beta}_t$ are obtained by regressing $\bar{y}$ on $\bar{z}$, where:

$$\bar{y} = [y_1, (1-\bar{\alpha}L)y_2, \ldots, (1-\bar{\alpha}L)y_T] \tag{5}$$

$$\bar{z} = [z_1, (1-\bar{\alpha}L)z_2, \ldots, (1-\bar{\alpha}L)z_T]\ . \tag{6}$$

Elliott, Rothenberg and Stock argue that fixing $\bar{c} = -7$ in the drift model, and $\bar{c} = -13.5$ in the linear trend model, used in (7) and (8), the test IS within 0.01 of the power envelope:

$$z_t = (1, t)' \tag{7}$$

$$\bar{\alpha} = 1 + \frac{\bar{c}}{T}\ . \tag{8}$$

4.2. Testing for unit roots in the presence of structural breaks

Perron (1989) has carried out tests of the unit root hypothesis against the alternative hypothesis of trend stationarity with a break in the trend. The two breakpoints included were 1929 (the Great Crash), and 1973 (the Oil Shock). Perron analysed the Nelson and Plosser (1982) macroeconomic data and quarterly post-war GNP series. His results rejected the unit root null hypothesis for most time series. Three models were considered:

$$y_t = \alpha_1 + \beta_1 t + (\alpha_2 - \alpha_1)DU_t + e_t, \quad t=1,2,\ldots T \tag{9}$$

$$y_t = \alpha_1 + \beta_1 t + (\beta_2 - \beta_1)DT_t + e_t, \quad t=1,2,\ldots T \tag{10}$$

$$y_t = \alpha_1 + \beta_1 t + (\alpha_2 - \alpha_1)DU_t + (\beta_2 - \beta_1)DT_t + e_t, \quad t=1,2,\ldots T \tag{11}$$

where, $DT_t = \begin{cases} t & if \quad t > TB \\ 0 & else \end{cases}$

and $DU_t = \begin{cases} 1 & if \quad t > TB \\ 0 & else \end{cases}$.

Equation (9) considers an exogenous break in the intercept, (10) an exogenous break in the trend, and (11) considers a break in both trend and intercept. To account for the possible serial autocorrelation, lagged values of the dependent variables can be included in the regression. The problem with the Perron test is that the breakpoint must be known *a priori* which is a seriously restrictive assumption. Zivot and Andrews (1992) modified the Perron test, to endogenously search for the breakpoints. That is achieved by computing the t-statistics for all breakpoints, then choosing the breakpoint selected by the smallest t-statistic, that being the least favourable one for the null hypothesis.

3.3. Cointegration analysis

The two most widely used cointegration tests are the Engle-Granger (Engle and Granger, 1987) two-step method and Johansen's multivariate approach (Johansen, 1988). Engle and Granger base their analysis on testing the stationarity of the error term in the cointegrating relationship. An OLS regression is run with the studied variables, and the residuals are tested for unit roots. If the null of non-stationarity can be rejected the variables are considered to be cointegrated.

The Johansen testing procedure has the advantage that allows for the existence of more than one cointegrating relationship (vector) and the speed of adjustment towards the long-term equilibrium is easily computed. The 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 the Vector Error Correction Model (VECM) of the form:

$$\Delta \mathbf{Z}_t = \mathbf{\Gamma_1}\Delta Z_{t-1} + \ldots + \mathbf{\Gamma_{k-1}}\Delta Z_{t-k+1} + \mathbf{\Pi} Z_{t-k} + \mathbf{\Psi D} + \mathbf{u_t} \qquad (12)$$

where Z_t = [$P^R{}_t$, $P^P{}_t$]' a (2 x 1) vector containing the farm and retail price, both I(1), $\mathbf{\Gamma_1}$,.... $\mathbf{\Gamma_{k+1}}$ are (2x2) vectors of the short run parameters, $\mathbf{\Pi}$ is (2x2) matrix of the long-run parameters, $\mathbf{\Psi}$ is a (2x11) matrix of parameters , **D** are 11 centred seasonal dummies and $\mathbf{u_t}$ is the white noise stochastic term.

$\mathbf{\Pi} = \boldsymbol{\alpha\beta}$`, where matrix $\boldsymbol{\alpha}$ represents the speed of adjustment to disequilibrium and $\boldsymbol{\beta}$ is a matrix which represents up to (n - 1) cointegrating relationships between the non-stationary variables. There are several realistically possible models in (12) depending on the intercepts and linear trends. Following Harris (1995) these models defined as models 2-4, are: (M2) the intercept is restricted to the cointegration space ; (M3) unrestricted intercept no trends - the

intercept in the cointegration space combines 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.
Because usually is not known *a priori* which model to apply, the Pantula principle (Harris 1995) is used to simultaneously test for the model and the cointegration rank.

4. Data and results

The theoretical model developed by Saghaian et al. (2002) serves as a guide for our empirical work. This model supposes a small open economy which is an appropriate assumption for Hungary. Monthly time series of an agricultural variable, the log of producer price index (lnPPI), the log of industrial producer price index (lnIPI), the log of Euro/Hungarian Forint exchange rate and the log of the seasonally adjusted money supply (M1A) were used. The dataset presented on figures 1 and 2, covers the January 1997 – August 2004 period, consisting of 92 observations. Data sources are the CSO-Central Statistical Office, and NBH – National Bank of Hungary.

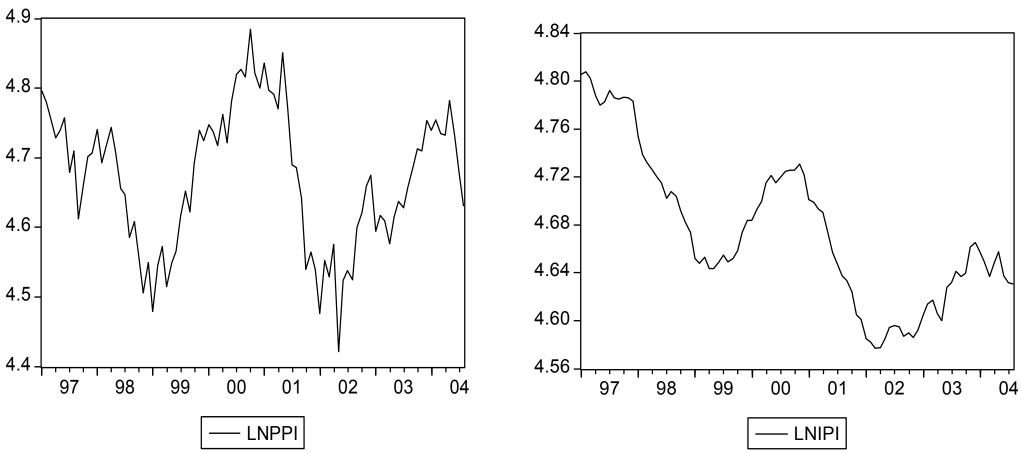

Figure 1. The logs of agricultural producer and industrial producer price indexes

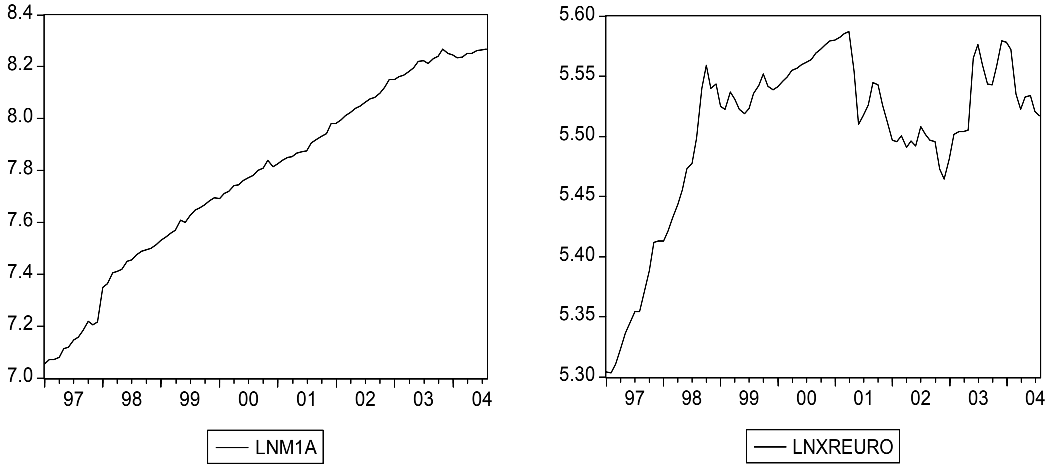

Figure 2. The logs of seasonally adjusted money supply and exchange rate

4.1. Stationarity and integration tests

First, the Elliott, Rothenberg, Stock (1996) DF-GLS unit root test, with and without a linear trend is performed. The results are presented in Table 1.

Table 1. DF-GLS unit root tests

Variables	Specification	Lags	Test statistic
lnIPI	constant	5	- 0.904
	constant and trend	5	- 2.997
lnPPI	constant	3	- 1.722
	constant and trend	3	- 2.349
lnM1A	constant	5	0.366
	constant and trend	5	- 0.697
lnXREURO	constant	2	-0.264
	constant and trend	2	- 0.931

The critical values for 0.95 (0.99) confidence levels with constant are -1.944 (-2.592), with constant and trend are -3.074 (-3.633). The Akaike Information Criteria was used to determine the lag length.

None of the tests statistics is significant, all the variables appears to be integrated. To ensure that all series are I(1), and not integrated of a higher order, the first differences are tested using the DF-GLS unit root tests in table 2. Because there is no evidence of a linear trend in the first difference representation of the variables, we conduct the second order unit root tests on the model with a drift only. The unit root null hypothesis is rejected at conventional levels for all series in first differences.

Table 2. DF-GLS unit root tests on the first differences of series

Variables	Specification	Lags	Test statistic
ΔlnIPI	constant	4	- 1.986
ΔlnPPI	constant	2	- 3.680
ΔlnM1A	constant	1	- 6.633
ΔlnXREURO	constant	1	- 6.753

The critical values for 0.95 (0.99) confidence levels with constant are -1.944 (-2.592), with constant and trend are -3.074 (-3.633). The Akaike Information Criteria was used to determine the lag length.

Because the graphical analysis suggests the possibility of structural breaks, that are not accounted for by the DF-GLS tests, the Zivot-Andrews unit root test in the presence of structural breaks is used to double-check the series (table 3).

Table 3. Andrews – Zivot unit root tests

Variable	Structural break	Lags	Possible break date	Test statistic
lnPPI	intercept only	0	2001:06	- 3.783
	trend only	0	2000:09	- 2.506
	both	0	2001:06	- 3.825
lnIPI	intercept only	1	2003:06	- 2.557
	trend only	1	1998:03	- 2.413
	both	1	2001:05	- 2.865
lnXREURO	intercept only	2	2001:05	- 3.717
	trend only	2	1998:09	- 3.278
	both	2	1998:08	- 3.263
lnM1A	intercept only	0	2003:07	- 2.208
	trend only	0	1998:03	- 3.655
	both	0	1998:08	- 2.876

The critical values for 0.95 (0.99) confidence levels in case of a structural break in the intercept only, are –4.80 (-5.34), trend only –4.42 (-4.93), and both -5.08 (-5.57). The Schwarz Bayesian criterion was used to determine the lag length.

Most hypothesised structural breaks appear after 2000, but none of the t-statistics associated with the possible structural break dates are statistically significant at .95% confidence level. Thus the results of the Zivot-Andrews test reinforces the DF-GLS test result, thus we consider all series being integrated of order 1.

4.2. Cointegration tests

First, the VECM lag length was selected. The various lag length criteria suggested different lag lengths, ranging between 1 (Schwarz-Bayesian Criterion), and 12 (Akaike Information Criterion). 5 lags in the VAR model were considered enough to result uncorrelated residuals, the Final Prediction Error and LR statistic also selecting the same lag length. The *Pantula principle* selected model 4, where there is a trend restricted to the cointegration space. The cointegration test results are presented in table 4 and 5.

Table 4. Johansen cointegration test results – trace statistics

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5% Critical Value	1% Critical Value
None	0.326620	90.97482	62.99	70.05
At most 1	0.279446	56.96656	42.44	48.45
At most 2	0.220366	28.78137	25.32	30.45
At most 3	0.082164	7.373335	12.25	16.26

Table 5. Johansen cointegration test results – max Eigen statistics

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	5% Critical Value	1% Critical Value
None	0.326620	34.00826	31.46	36.65
At most 1	0.279446	28.18519	25.54	30.34
At most 2	0.220366	21.40804	18.96	23.65
At most 3	0.082164	7.373335	12.25	16.26

The trace statistics selects 3 cointegration vectors at 5% level and 2 cointegration vectors at 1%, level, whilst the maximum Eigen statistic selects 3 cointegration equations at 5% level.

We conclude 3 cointegration vectors at 5% level of significance. The normalised cointegration vectors are presented in table 6.

Table 6. Normalized cointegrating coefficients

lnPPI	lnIPI	lnXREURO	lnM1A	TREND
1.000000	0.000000	0.000000	0.100722	0.000237
			(0.40240)[a]	(0.00539)
0.000000	1.000000	0.000000	0.432500	-0.003577
			(0.12665)	(0.00170)
0.000000	0.000000	1.000000	-0.648281	0.008627
			(0.12772)	(0.00171)

[a] standard errors in parentheses

The money slope coefficients are rather surprisingly negative for the industrial and agricultural prices and positive for the exchange rate equation, not being statistically significant in the agricultural price equation. The linear trend is significant in the industrial prices and exchange rate equations, but not in the agricultural prices equation.

The money neutrality hypothesis expects the coefficients associated with the money supply (lnM1A) to be close to one (i.e. the long run increase in the agricultural, industrial and services prices to be unit proportional with the increase in the money supply). The lnM1A coefficients with respect to the prices are 0.100, 0.432, -0.648, not supporting the money neutrality hypothesis.

4.3. VECM model

Because the variables proved to be cointegrated, a Vector Error Correction Model is appropriate to simultaneously depict the long and short run evolution of the system. The residuals of the long run cointegrating equations are used to construct the VECM in table 7.

Table 7. Vector error correction model coefficients[a] and diagnostic tests

Cointegrating Equations	CointEq1	CointEq2	CointEq3	
$lnPPI_{t-1}$	1.000000	0.000000	0.000000	
$lnIPI_{t-1}$	0.000000	1.000000	0.000000	
$lnXREURO_{t-1}$	0.000000	0.000000	1.000000	
$lnM1A_{t-1}$	0.100722 [0.24636][b]	0.432500 [3.36114]	-0.648281 [-4.99590]	
TREND	0.000237 [0.04329]	-0.003577 [-2.07654]	0.008627 [4.96603]	
C	-5.465342	-7.872801	-0.869331	
Error Correction:	$\Delta lnPPI_t$	$\Delta lnIPI_t$	$\Delta lnXREURO_t$	$\Delta lnM1A_t$
Coint.Eq1	*-0.479967* [-3.09890]	0.013393 [0.43138]	-0.086666 [-1.96655]	0.047808 [0.69695]
CointEq2	0.906589 [1.72319]	*-0.121941* [-1.15627]	0.435903 [2.91188]	-0.180480 [-0.77456]
CointEq3	0.093395 [0.41178]	-0.020625 [-0.45366]	*-0.298322* [-4.62266]	-0.171357 [-1.70589]
$\Delta lnPPI_{t-1}$	-0.113174 [-0.72505]	-0.025751 [-0.82301]	0.037446 [0.84312]	0.042338 [0.61243]
$\Delta lnPPI_{t-2}$	0.200775 [1.46324]	-0.023857 [-0.86737]	-0.028477 [-0.72940]	0.053514 [0.88060]
$\Delta lnPPI_{t-3}$	0.331507 [2.60738]	0.017568 [0.68931]	0.018086 [0.49994]	0.050697 [0.90032]
$\Delta lnIPI_{t-1}$	0.011583 [0.01520]	0.448452 [2.93546]	0.010429 [0.04809]	-0.409869 [-1.21430]
$\Delta lnIPI_{t-2}$	-0.113043 [-0.16562]	-0.003211 [-0.02347]	-0.468083 [-2.41025]	-0.293916 [-0.97231]
$\Delta lnIPI_{t-3}$	-0.036470 [-0.05327]	0.019099 [0.13916]	0.105485 [0.54147]	0.299201 [0.98670]
$\Delta lnXREURO_{t-1}$	-0.521604 [-1.18082]	-0.172611 [-1.94938]	0.162462 [1.29258]	0.257264 [1.31500]

$\Delta lnXREURO_{t-2}$	-0.082630	0.136039	-0.094994	0.045497
	[-0.17979]	[1.47662]	[-0.72641]	[0.22352]
$\Delta lnXREURO_{t-3}$	-0.350875	-0.041217	-0.320742	0.045060
	[-0.76772]	[-0.44989]	[-2.46641]	[0.22261]
$\Delta lnM1A_{t-1}$	-0.344271	0.001166	-0.266600	-0.362669
	[-0.81551]	[0.01378]	[-2.21948]	[-1.93974]
$\Delta lnM1A_{t-2}$	-0.040444	0.024642	-0.226520	-0.223882
	[-0.09884]	[0.30042]	[-1.94553]	[-1.23536]
$\Delta lnM1A_{t-3}$	0.436417	-0.072496	-0.222130	-0.084588
	[1.14138]	[-0.94586]	[-2.04173]	[-0.49951]
C	0.008236	-0.000118	0.014232	0.023670
	[0.43235]	[-0.03084]	[2.62579]	[2.80560]
R^2	0.509773	0.522253	0.566863	0.292089
Adj. R^2	0.327914	0.345025	0.406183	0.029476
Akaike criterion	-3.365201	-6.579540	-5.878990	-4.994069
Schwarz criterion	-2.680267	-5.894606	-5.194055	-4.309135
Jarque-Bera	4.858*	3.85	5.903**	100.116***

[a] because of space limitations, coefficients up to the 3rd lag are shown only
[b] t-statistics in brackets
Note: ***1% significance level, **5% significance level, ***10% significance level

The coefficients of the three cointegration equations in the VECM, called the speeds of adjustment (**α** in equation 12), measure how quickly the system returns to its long run equilibrium after a temporary shock. More exactly, if say, the agricultural prices are overshooting their long run equilibrium path, then the associated **α** value must be negative, implying that prices must fall in order to re-establish the long run equilibrium between money supply and prices. By considering one flexible (agriculture and exchange rate) and one sticky (industry) sector, we would expect to have larger (in absolute value) **α** parameters associated with flexible sector prices than with the sticky sector prices (Shagaian et al. 2002). The speeds of adjustment to the long run equilibrium of the agricultural, industrial prices and exchange rate are -0.4799, -0.1219, -0.2983 (table 7, in Italic), all negative as expected and significant, except industrial prices. More, the values associated with flexible sector prices are bigger (in absolute values) than the one associated with the industrial prices, suggesting a faster adjustment of the flexible sector, result also consistent with the literature. None of the error correction terms seem to be significant in the industrial price equation, suggesting exogeneity (industrial prices would not adjust after a shock to the system), but a joint zero restriction of the speed of adjustment vector is rejected ($\chi^2(3) = 9.807$, $p = 0.02$).

The coefficients of determination are similar to those obtain by other studies ranging between 0.29 and 0.57, thus the model explains a relatively high percent of change in the macroeconomic variables. The Jarque-Bera statistics reject the normality null at 10% for 3 equations. However, non-normality – implies that the test results must be interpreted with care, although asymptotic results do hold for a wider class of distributions (von Cramon-Taubadel, 1998).

Table 8. Residual serial autocorrelation LM and LB tests

Lags	LM-Stat	Prob.[a]	Lags	LM-Stat	Prob.
1	19.18801	0.2590	7	8.600210	0.9290
2	16.41018	0.4247	8	15.34749	0.4994
3	11.53637	0.7752	9	21.08346	0.1753
4	16.56960	0.4140	10	10.37361	0.8464
5	21.45633	0.1616	11	11.87551	0.7525
6	20.28460	0.2077	12	21.57624	0.1574
Ljung-Box statistic (21)	$\chi^2(244) = 288.472$ ($p = 0.03$)				

[a] Probabilities from chi-square with 16 df.

Multivariate LM tests for serial autocorrelation do not reject the no-autocorrelation null hypothesis for up to the 12[th] order, but the no-autocorrelation in the first 21 observations null is rejected.

5. Conclusions

In this research, a theoretical model developed by Shagaian et al. (2002) was employed for a small, open economy. As most post-communist economies, Hungary experienced numerous monetary shocks during the transition period, many of them due to the less developed monetary instruments and ad-hoc measures. Empirical evidence is presented that these shocks quickly found their way into the agricultural sector causing significant though largely unmapped effects. The existence of three cointegration vectors amongst the Hungarian agricultural prices, industrial prices, exchange rate, and money supply, proves the existence of a long-run equilibrium relationship between the variables. It follows, that shocks to macroeconomic variables find their way onto the agricultural sector. After identifying the cointegrating equations and examining the slope coefficient of the money supply, we found that the money neutrality hypothesis doesn't hold for Hungary. In accordance with the theoretical model mentioned above, we found evidence that agricultural prices adjust faster to monetary shocks than industrial prices do. The other flexible sector considered (the exchange rate) also adjusts faster to temporary shocks than the sticky, industrial sector. Thus, if a monetary shock occurs, the flexible sectors will have to bear the burden of adjustment, reducing the financial viability of the Hungarian farmers.

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