Monetary Impacts and Overshooting of Agricultural Prices in a Transition Economy

The Case of Slovenia

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Abstract: The 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, and Bakucs and Fertő (2005) who tested the overshooting hypothesis on the Hungarian agricultural prices.

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 Slovenia 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

Slovenian 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 Slovenian 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):

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

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

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

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

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

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:

$$z_t = (1, t)' \quad (7)$$

$$\bar{\alpha} = 1 + \frac{\bar{c}}{T} \ . \quad (8)$$

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:

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

where Z_t = [PPI_t, IPI_t, XR_t, $M1_t$]' a (4 x 1) vector containing the four I(1) variables, $\mathbf{\Gamma_1}$,....$\mathbf{\Gamma_{k+1}}$ are vectors of the short run parameters, $\mathbf{\Pi}$ is matrix of the long-run parameters, $\mathbf{u_t}$ is the white noise stochastic term.

$\mathbf{\Pi = \alpha\beta}$`, 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 five possible models in (12) depending on the intercepts and linear trends. Following Harris and Sollis (2003) these models defined as models 1-5, are: (M1) no intercept or trend is included; (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; (M5) allowing for quadratic trends in Z_t.

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 Slovenia. Monthly time series of an agricultural variable, the log of producer price index (PPI), the log of industrial producer price index (IPI), the log of Euro/Slovenian Tollar exchange rate and the log of the money supply (M1) were used. The summary statistics of the used variables are presented in table 1. The dataset covering the January 1996 – July 2005 period, consisting of 115 observations are presented on figures 1 and 2. Data sources are the Slovenian Statistical Office, and the Bank of Slovenia.

Table 1. Summary statistics

Variable	Mean	Standard deviation	Minimum	Maximum
Agricultural Producer Price Index, PPI	100.77	12.67	120.3	76.09
Industrial Producer Price Index, IPI	103.1	14.91	126	79.2
Slovenian Tollar/ Euro Ex. Rate, XR	207.71	24.31	239.82	163.55
Money Supply, M1 (mil SIT)	867566.9	323388.8	1469809	403276

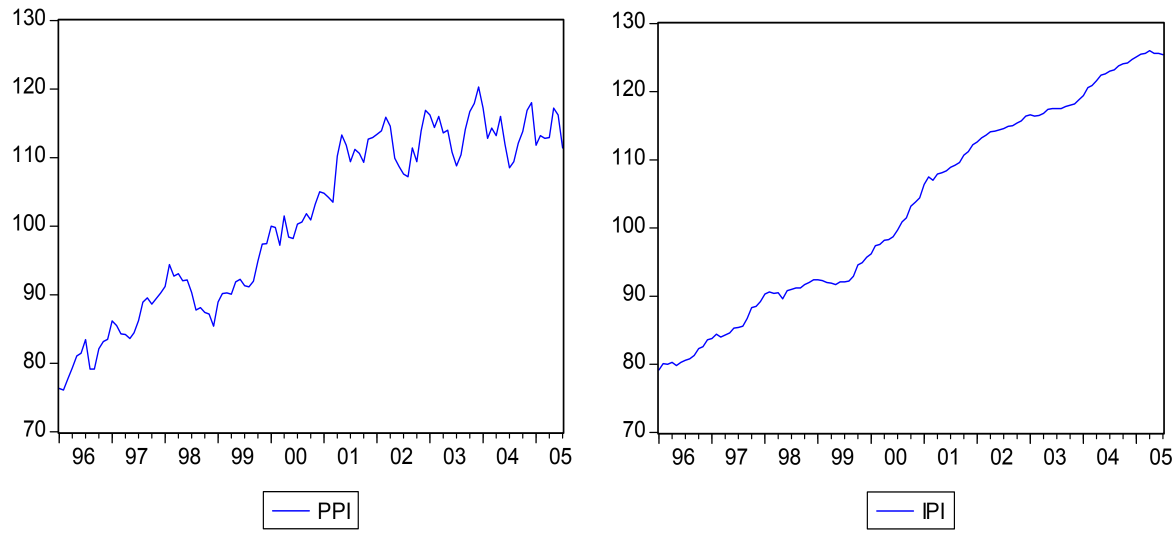

Figure 1. The agricultural producer and industrial producer price indexes

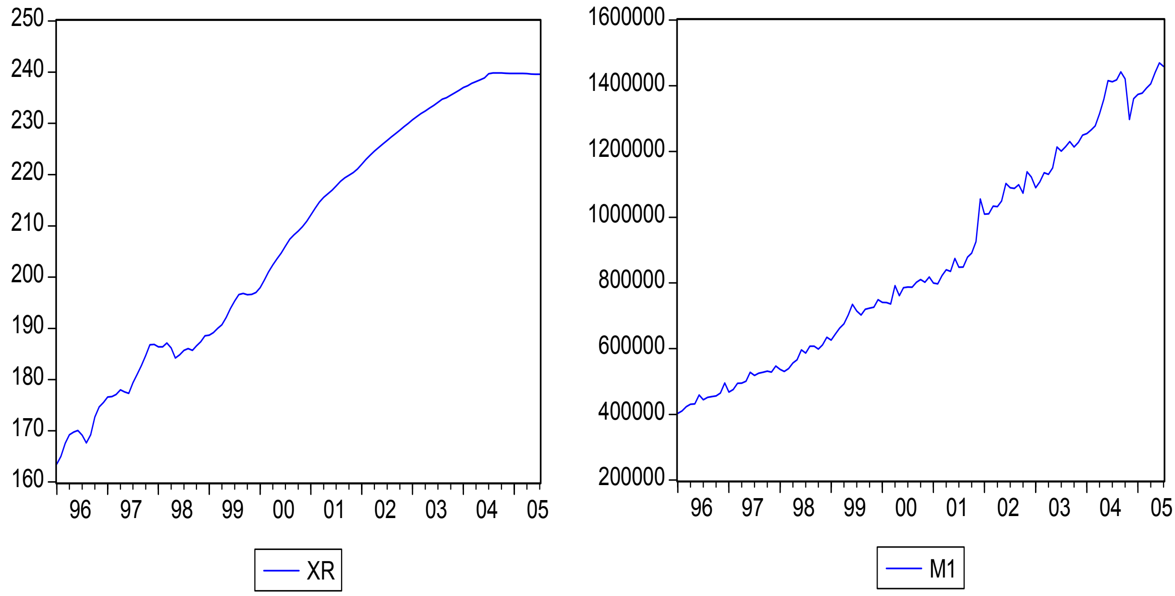

Figure 2. The exchange rate (SIT/Euro) and money supply (mil SIT)

4.1. Stationarity and integration tests

First, the Augmented Dickey-Fuller unit root tests (not shown here) with and without a trend are performed. Test results suggest all series are integrated of order one, I(1). Second, the more up-to-date Elliott, Rothenberg, Stock (1996) DF-GLS unit root test, with and without a linear trend is run. The results are presented in the first part of Table 2. None of the tests statistics is significant, all the variables appears to be integrated. To check whether all series are I(1) or integrated of a higher order, the first differences[1] are tested using the DF-GLS unit root tests. The results are presented in the second part of table 2. With or without a trend, the first difference of the industrial prices seems to be integrated of a higher order than one. The first difference of the exchange rate with constant does not reject the unit root null either. At this point, two issues need to be mentioned. First, the often poor size and power properties of the unit root tests may lead to unbalanced results. Second, it is possible that cointegration exists when there is a mix of variables integrated of different order as the variables integrated of order 2 can first cointegrate down to I(1), than cointegrate with the rest of the variables resulting stationary residuals (Harris and Sollis, 2003, pp.112).

Therefore we test the cointegration between the four variables, than analyse the stationarity properties of the resulting residuals.

[1] the graphs of the first difference series are presented in the appendix

Table 2. DF-GLS unit root tests on the variables

Variables	Specification	Lags	Test statistic
PPI	constant	0	0.245
	constant and trend	0	- 1.798
IPI	constant	10	0.235
	constant and trend	0	- 0.779
XR	constant	1	1.485
	constant and trend	1	- 0.841
M1	constant	6	1.606
	constant and trend	6	- 2.942
First differences			
ΔPPI	constant	2	- 6.34
	constant and trend	0	- 9.104
ΔIPI	constant	9	- 1.252
	constant and trend	9	- 2.148
ΔXR	constant	0	- 1.933
	constant and trend	1	- 3.854
ΔM1	constant	5	- 3.402
	constant and trend	5	- 4.096

The critical values for 0.95 (0.99) confidence levels with constant are -1.943 (-2.585), with constant and trend are -3.015 (-3.562). The Schwarz Bayesian Criteria was used to determine the lag length.

4.2. Cointegration tests

First, the VECM lag length was selected. Three of the five usual lag length criteria, the LR test statistic, the final prediction error criteria (FPE) and the Akaike Information Criterion (AIC) suggested 4 lags, whilst the Schwarz Bayesian Criterion (SBC) and the Hannan-Quinn (HQ) suggested 2 lags. 4 lags in the VAR model were considered enough to result uncorrelated residuals. The number of cointegrating vectors depends on the model specification chosen (M1 - M5), however at least one (trace statistic) or two (Max- Eigen statistic) cointegration vectors were found at 5 % significance level irrespectively of the model specification. Specification M5 (quadratic trends) was found to maximise the log likelihood function and also to minimise the AIC criteria. It might be difficult to argue in favour of quadratic trends in economic processes, it should be noted however that all specifications are nested in M5, therefore the misspecification bias is minimised by using this specification. The cointegration test results are presented in table 3.

Table 3. Johansen cointegration test results – trace statistics and max Eigen statistics

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	5% Critical Value	1% Critical Value
None	0.296087	84.6008	55.2457	0.000
At most 1	0.200094	44.2243	35.0109	0.004
At most 2	0.145316	18.5492	18.3977	0.047
At most 3	0.004265	0.4915	3.8414	0.483
Hypothesized No. of CE(s)	**Eigenvalue**	**Max-Eigen Statistic**	**5% Critical Value**	**1% Critical Value**
None	0.296087	40.3765	30.8150	0.002
At most 1	0.200094	25.6750	24.2520	0.032
At most 2	0.145316	18.0576	17.1476	0.036
At most 3	0.004265	0.4915	3.8416	0.483

We conclude 3 cointegration vectors at 5% level of significance. Because of the ambiguous unit root test results, the three cointegration residuals (fig. 3) are tested for unit roots. The test results (table 4) reject the unit root null hypothesis for all three residuals at 1% level of significance. Table 5 presents the normalised cointegration vectors.

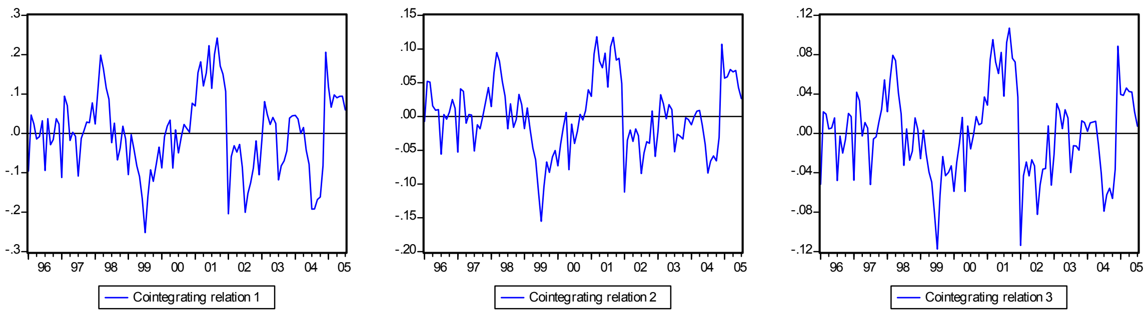

Figure 3. The cointegration residuals

Table 4. DF – GLS unit root tests on the cointegration residuals

Variables	Specification	Lags	Test statistic
cointeg1	constant[2]	0	- 3.609
cointeg2	constant	0	- 4.193
cointeg3	constant	0	- 3.606

The critical values for 0.95 (0.99) confidence levels with constant are -1.943 (-2.585). The Schwarz Bayesian Criteria was used to determine the lag length.

Table 5. Normalized cointegrating coefficients

PPI	IPI	XR	M1
1.000000	0.000000	0.000000	- 2.587 (0.6137)
0.000000	1.000000	0.000000	- 1.382 (0.3587)
0.000000	0.000000	1.000000	- 1.258 (0.2723)

[a] standard errors in parentheses

The empirical long-run relationships between the producer price and money supply, industrial price and money supply, exchange rate and money supply are in line with our expectations. The money slope coefficients are all negative and significant, consistent with economic theory that expansionary monetary policy positively affects prices. The money neutrality hypothesis expects the coefficients associated with the money supply (M1) 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). One percent increase in money supply results in 2.587%, 1.382% and 1.258% increase in the agricultural producer prices, industrial prices and exchange rate respectively, not supporting the money neutrality hypothesis.

To test the long run neutrality hypothesis of the individual long-run relationships, restrictions are imposed on the M1 coefficients. The restriction is rejected for the producer price ($p = 0.042$), but couldn't be rejected for the industrial prices and exchange rate equations ($p = 0.567$ and $p = 0.838$ respectively). However, it appears that money supply in Slovenia is not neutral.

[2] Although the graphical inspection does not suggest, a constant is needed to perform the DF-GLS test. The ADF tests were also run with and without a constant and rejected the unit root null 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 6.

Table 6. Vector error correction model coefficients and diagnostic tests

Cointegrating Equations	CointEq1	CointEq2	CointEq3	
PPI_{t-1}	1.000000	0.000000	0.000000	
IPI_{t-1}	0.000000	1.000000	0.000000	
XR_{t-1}	0.000000	0.000000	1.000000	
$M1_{t-1}$	- 2.587346 [- 4.17340][a]	- 1.382735 [- 3.81612	- 1.258575 [- 4.57480	
TREND	0.026006	0.011515	0.010895	
C	28.45727	13.23992	10.89726	
Error Correction:	ΔPPI_t	ΔIPI_t	ΔXR_t	$\Delta M1_t$
Coint.Eq1	*- 0.580789* [- 5.18889]	0.054858 [2.06287]	0.029633 [1.53241]	0.077715 [0.57599]
CointEq2	0.062575 [0.43088]	*- 0.122915* [- 3.56234	0.051831 [2.06581	- 0.164856 [- 0.94172
CointEq3	1.194650 [4.16526]	0.015664 [0.22987]	*- 0.142944* [- 2.88474]	0.275233 [0.79609]
ΔPPI_{t-1}	0.303185 [2.79528]	- 0.024980 [- 0.96934]	- 0.029097 [- 1.55278]	- 0.185611 [- 1.41964
ΔPPI_{t-2}	0.119705 [1.14835]	- 0.029317 [- 1.18373]	- 0.008751 [- 0.48593]	- 0.119826 [- 0.95361]
ΔPPI_{t-3}	0.154899 [1.53327]	- 0.049732 [- 2.07195]	- 0.008984 [- 0.51471]	- 0.149138 [- 1.22466]
ΔIPI_{t-1}	0.144382 [0.37347]	0.069826 [0.76021]	0.099311 [1.48689]	-1.549886 [- 3.32582]
ΔIPI_{t-2}	- 0.231426 [- 0.56445]	0.186723 [1.91683]	0.085938 [1.21322]	1.209131 [2.44649]
ΔIPI_{t-3}	0.956857 [2.24765]	0.115513 [1.14205]	- 0.063974 [- 0.86981]	- 1.524581 [- 2.97091]
ΔXR_{t-1}	- 0.001207 [- 0.00212]	0.109770 [0.81085]	0.708433 [7.19648]	0.068735 [0.10007]
ΔXR_{t-2}	- 0.942684 [- 1.47268]	- 0.065537 [- 0.43093]	- 0.250921 [- 2.26892]	0.838802 [1.08707]
ΔXR_{t-3}	- 0.124408 [- 0.22610]	0.187454 [1.43391]	0.141029 [1.48355]	- 0.293354 [- 0.44228]

$\Delta M1_{t-1}$	0.034907 [0.38140]	- 0.019553 [-0.89919]	- 0.005576 [- 0.35266]	0.005107 [0.04629]
$\Delta M1_{t-2}$	0.063998 [0.78392]	- 0.019327 [- 0.99643]	- 0.027254 [- 1.93230]	- 0.081625 [- 0.82944]
$\Delta M1_{t-3}$	0.008671 [0.11877]	- 0.032453 [- 1.87107]	0.003567 [0.28284]	0.224166 [2.54733]
C	0.006039 [0.82540]	0.003805 [2.18909]	0.002886 [2.28305]	0.021765 [2.46797]
Trend	- 7.26E-05 [- 1.16893]	- 1.22E-05 [- 0.82645]	- 1.88E-05 [- 1.74961]	- 6.16E-05 [- 0.82309]
Adj. R^2	0.208738	0.247946	0.520961	0.428191
Log Likelihood	300.7065	465.9863	502.6250	279.2203
Akaike criterion	- 4.934027	- 7.808457	- 8.445653	- 4.560352
Schwarz criterion	- 4.528254	- 7.402685	- 8.039880	- 4.154580
Jarque-Bera	5.052573*	2.658755	1.141660	41.58510***

[a] 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 -0.5807, -0.1229, -0.1429 (table 6, in Italic), all negative as expected and significant. 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.

Because of the difficulty to interpret VAR coefficient estimates, it is common to employ impulse response functions to graphically depict the influence of a shock upon the VAR variables. The generalised impulse response functions of Pesaran and Shin (1998) were used to simulate the responses of the agricultural producer prices, industrial prices and the exchange rate upon a one standard deviation shock in the money supply (figure 4).

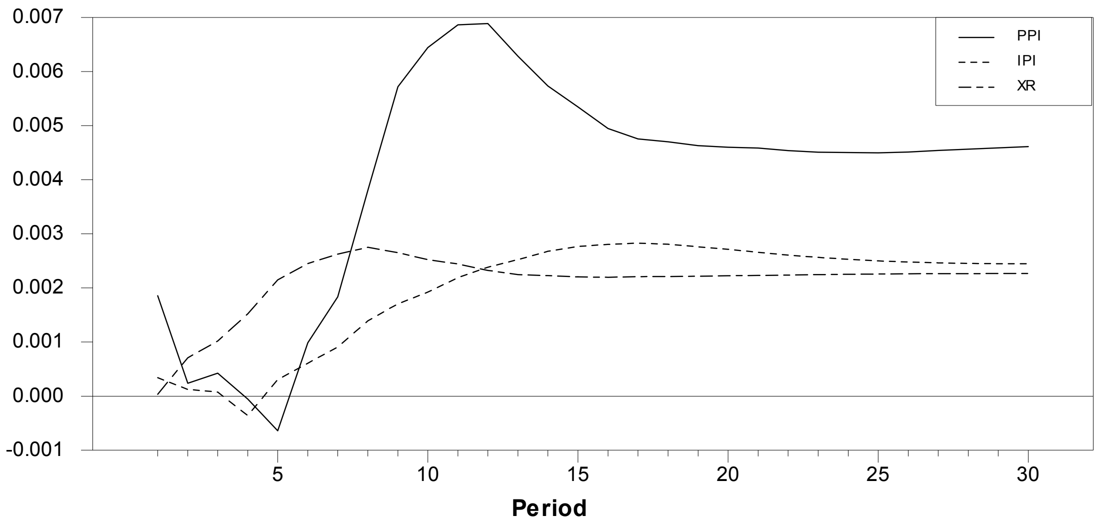

Figure 4. Impulse response to one standard deviation of money supply shock

The impulse response analysis reinforces the previous results. An exogenous shock to the money supply has a significant and volatile effect on the three price variables. First, both the agricultural and industrial prices undershoot their long-run path, the negative jump affecting agriculture being twice as large as the one for industrial prices. Industrial prices recover in 4, agricultural prices in 6 months after the initial shock, than overshoot their long-run equilibrium. Supporting both the theoretical model and previous results, the agricultural prices experience the largest overshooting (twice as much as exchange rates or industrial prices). The monetary shock has a persistent effect on all three prices, they stabilise around a new equilibrium path in approximately 17 – 20 month after the original shock occurred.

A different tool to analyse the VAR results is the forecast error variance decomposition (table 7).

Table 7. Variance decomposition for PPI

Period	PPI	IPI	XR	M1
1	92.40038	0.227640	6.437224	0.934760
2	85.26632	0.549112	13.60832	0.576246
3	78.42139	0.736631	20.32645	0.515534
4	66.86413	5.248300	27.45670	0.430870
5	55.83511	8.297101	35.46780	0.399989
6	47.70034	9.832854	42.04403	0.422774
12	30.61714	9.839003	50.93145	8.612409
24	22.81957	9.038791	54.91346	13.22818

Variance decomposition separates the variation in an endogenous variable into the component shocks to the VAR. Thus, the variance decomposition provides information about the relative importance of each random innovation in affecting the variables in the VAR.
Interesting and perhaps intriguing results were obtained. On a 12 month horizon only 30% of the variation in the industrial prices is explained by its own shock (e.g. changes in the supply demand conditions), and only 8.6% of the variation is due to money supply factors. Exchange rate however, seems to play an unusually important role (50%) in the explaining the expected variance in the agricultural prices. On a 2 year horizon, the effect of the own variation further diminishes (23%), whilst the percentage of variation explained by money supply and exchange rate variation increases (13.2% and 55% respectively).
What could explain the importance of the exchange rate in the expected variation of the agricultural prices? Agricultural imports in Slovenia amount to around 30%, exports to approximately 5% of the total agricultural output. One may argue that because Slovenia is a small, open economy, agricultural prices quickly adjust to the international prices through the exchange rate.
The coefficients of determination (lower part of table 6) are similar to those obtained by other studies, ranging between 0.12 and 0.44, 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 2 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	19.075	0.264	7	25.520	0.061
2	11.020	0.808	8	8.794	0.921
3	13.640	0.625	9	22.970	0.114
4	11.873	0.752	10	15.364	0.498
5	12.952	0.676	11	16.033	0.450
6	7.645	0.958	12	25.665	0.0589
Ljung-Box statistic (16)	Adj. Q-stat = 220.86 (p = 0.257)				

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

Multivariate LM tests for serial autocorrelation (table 8) do not reject the no-autocorrelation null hypothesis at 5 % for up to the 12th order, and the Ljung - Box statistic indicates there is no autocorrelation amongst the first 16 lags.

5. Conclusions

This paper has examined the overshooting hypothesis for the Slovenian agriculture employing a theoretical model developed by Shagaian et al. (2002). As most post-communist economies, Slovenia experienced numerous monetary shocks during the transition period, many of them due to the less developed monetary instruments and ad-hoc measures. Our results suggest 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 Slovenian 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 Slovenia. 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 Slovenian farmers.

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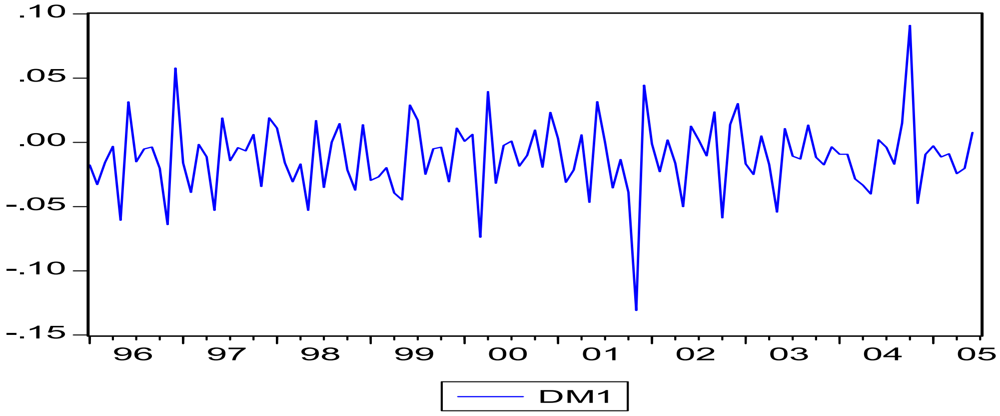
Appendix – first difference graphs of the series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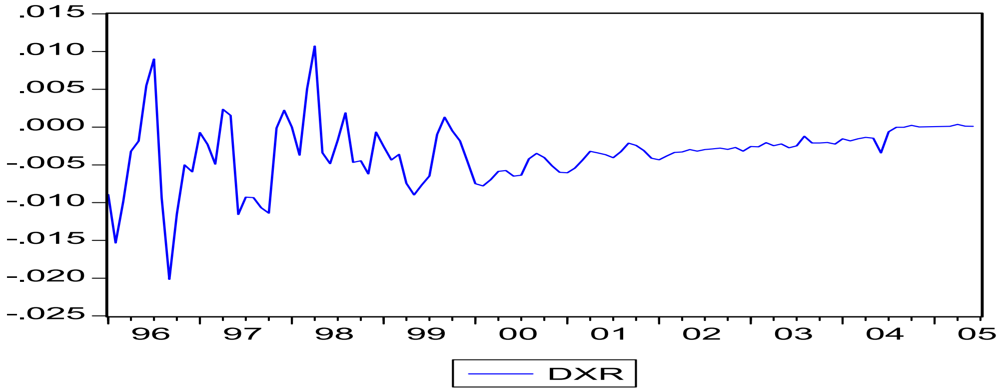
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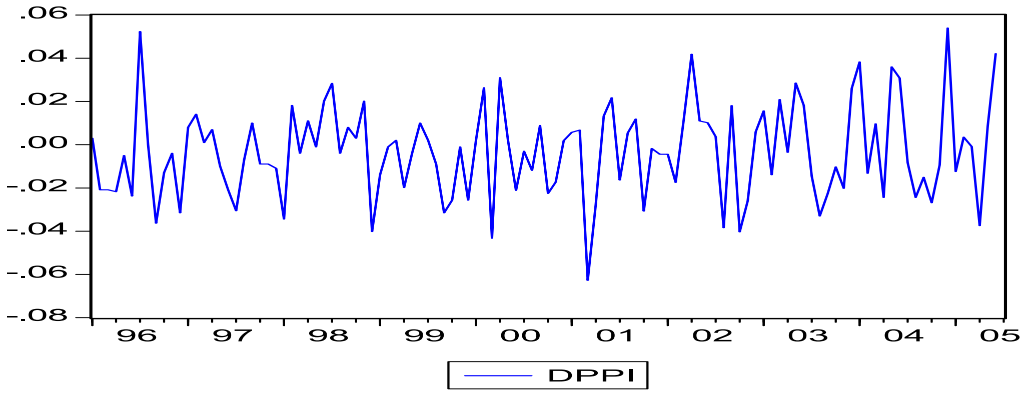
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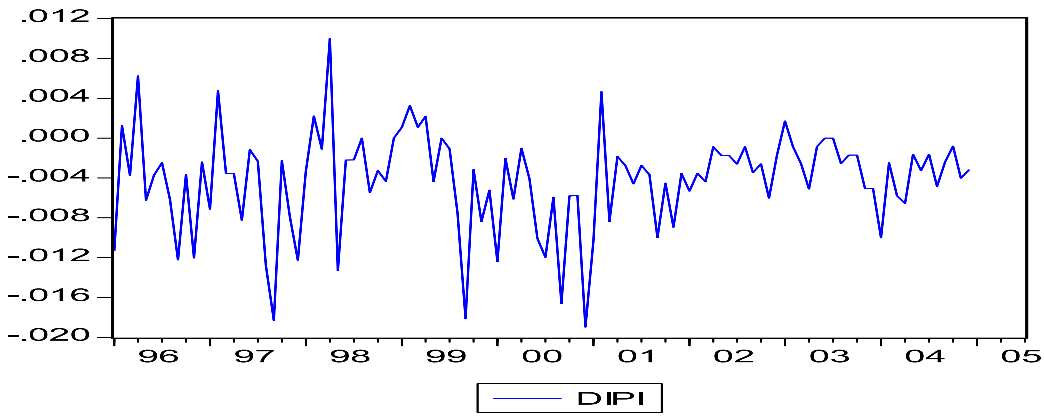
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