Study of Causal Dependencies among Prices in the Czech Wheat Commodity Market

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Abstract

Czech commodity market with wheat underwent significant changes since the beginning of the market economy. In 2004, opening of the EU market brought new business opportunities for Czech plant producers. As a result, nearly third of the country’s wheat production is currently realized in foreign markets. This paper focuses on exploration of Granger causality relationships among average wheat price of human consumption quality; animal feed quality and the average price of wheat in the world market (source: World Bank). Monthly time series from July 1993 to August 2013 (T=242) were subjected to stationarity and cointegration tests. Due to nonexistence of equilibrium relationships, the multivariate data were analyzed by Vector Autoregressive (VAR) model in first differences. Separate VAR models were fitted in periods before EU entry and after to verify assumptions of differential short-term dynamic relationships between the prices, as a result of country’s accession to EU. Further, orthogonal impulse-response function was prepared to quantify impact of innovation shocks in the source series upon the target series. Differences in relationships were found due to exposure of the Czech market to fluctuations in the world wheat price.

Keywords: EU accession; VAR model; Granger causality; impulse-response function; commodity market; wheat price

1. Introduction

Growing cereal crops has long and important tradition in the Czech Republic. Wheat production belongs to one of the most favored activities by Czech farmers, mainly because of strong tradition and professional support from the...
government and the agricultural research community. The area of agricultural land sowed with wheat varies around 0.8 mil. hectares every year, which constitutes about 33 % of all agricultural land. During years 1993–2011, the size of this land was relatively stable, despite the fact that land used for growing other agricultural crops steadily declined from approx. 2.4 mil. hectares in 1993 to about 1.6 mil. hectares in 2011. The reduction happened mainly as a result of downsizing in the entire agricultural sector.

The present Czech wheat market is strongly export oriented. Annual domestic production of wheat oscillates between 3.5 and 4.5 mil. metric tons, while the domestic consumption fluctuates between 3 and 3.5 mil. metric tons. Due to gradually dropping domestic demand, the resulting wheat surplus is often realized in foreign markets, which in recent years absorbed increasing amounts of unused wheat. This phenomenon became widespread especially after the country gained unlimited access to EU common market in May 2004.

In light of severe reduction that occurred in the Czech sugar beet and raw sugar market during 2004–2010, it was enormously encouraging for Czech wheat growers maintaining solid market position, that decreasing domestic consumption did not lead to cuts in wheat production, but instead, the wheat surplus was quickly redirected to foreign markets. Annual wheat exports from the Czech Republic amount between 0.8 and 1.6 mil. metric tons. Due to transportation costs, most wheat is exported to countries neighboring with the Czech Republic: Germany, Poland, Austria, Slovakia and Italy. Assuming that current demand for wheat in the world continues to increase in the following years, the volume of wheat exported from the Czech Republic to other EU countries is likely to grow. Good prospects also exist for moderate increase in wheat price leading to greater revenues for producers that could balance out poor earnings incurred from other activities, such as vegetables or animal production.

Syrovátka (2010) assumed that wheat price of food quality standard is determined primarily by supply and demand in the domestic market. On the other hand, Bečvářová (2004) stated that wheat price oscillations are determined mainly by foreign trade with this commodity. Inevitably, inclusion of the Czech Republic to the EU economic space was reflected in convergence of economic activity and increased similarity in most economic indicators. It is assumed, that free movement of agricultural commodities and products across borders of EU member states shall result in stronger correlation between commodity prices. At the same time, situation in the Czech domestic market also may have significant influence on the wheat price variation. Objective of this paper is to describe causal relationships between two average wheat prices in the Czech domestic market and the average wheat price in the world market. Also, attempt shall be made to investigate hypothesis, whether opening of the EU common market in 2004 had any significant impact upon the character of the dynamic relationships.

2. Material and methods

2.1. Material

Monthly time series of average wheat price of human consumption quality ($Y_{hc}$) and average wheat price of animal feed standard from the Czech market ($Y_{af}$) were retrieved from the commodity report supplied by the Czech Ministry of Agriculture (www.eagri.cz). The prices are expressed in Czech koruna (CZK) per metric ton. In addition, average price of wheat in the world market (USD per ton) was recovered from the World Bank. To account for variation in the currency exchange rate, the world wheat price in USD per ton was converted to Czech koruna ($Y_{wmn}$) using the average monthly CZK/USD exchange rate (source: the Czech National Bank). Monthly time series subjected to statistical analyses were dated from July 1993 to August 2013 ($T=242$). Fig. 1 depicts average price of wheat for human and animal consumption in the Czech market and the average world price per ton after transformation.
2.2. Statistical analyses

Numerous statistical analyses of time series require that multivariate time series in the system be firstly checked for weak (covariance) stationarity. Stationarity stipulates that expected value and variance be independent of time. Covariances may depend on time shift, but must be invariant to time or any other variable. Presence of unit root was tested by KPSS test (Kwiatkowski et al., 1992), which verifies the null hypothesis of stationarity against the alternative of unit root. It uses the test statistic

\[ \eta = \frac{T^{-2} \sum_{t=1}^{T} S_t^2}{\hat{\sigma}_q^2}. \]  

In this formula, \( S_t = \sum_{i=1}^{T} \hat{e}_i \) and \( \hat{\sigma}_q^2 \) is estimate of variance adjusted for autocorrelation pursuing procedure by Newey and West (1987). Since it appears that none of the investigated series moves around a linear trend, a positive constant was considered to be a sufficient deterministic component in the level auxiliary regression. We took into consideration notion of Syrovátka (2010), who stated that commodity price series would unlikely oscillate around zero. Results from KPSS test were compared to correlogram and outcome of the variance drop test.

Existence of long-run equilibrium relationships among nonstationary time series in the trivariate system was verified by EG cointegration test described in Engle and Granger (1987). Series with unit root were stationarized by applying at least one round of non-seasonal differencing with the purpose to eliminate the unit root. Number of differencing rounds was determined via procedure presented by Hyndman and Razbash (2014). It returns the minimum number of simple differencing required to pass the KPSS test at \( \alpha=0.05 \).

Multivariate system of price time series was transformed to reach stationarity and later analyzed by stationary vector autoregressive model VAR(\( p \)) with additive seasonal constants to account for regular seasonal movements following Lütkepohl (2005). Order of the VAR model was determined by likelihood ratio \( \chi^2 \)-test and also by search procedure that minimizes BIC information criterion. Seasonal fluctuations in the price data were assessed by exploring sample periodogram and then applying Fisher’s g-test of periodicity component with unknown frequency discussed by Brockwell and Davis (1991). Upon estimating the VAR model by OLS, the trivariate error term was checked for white noise properties by multivariate variant of the portmanteau test. In addition, univariate tests for autocorrelation (Ljung-Box), conditional heteroskedasticity (Lagrange multiplier test for ARCH(\( q \)) effect) and overall test of multivariate normality (Doornik-Hansen) were carried out. Significance of causal relationships between the series in Granger (1969) sense was tested with Error Sum of Squares Reduction \( F \)-test or it was inferred from \( t \)-tests of model coefficients.
Statistical analysis and preparation of plots were completed with R software version 3.0.2 (R Core Team, 2013) and Gretl 1.9.12 (http://gretl.sourceforge.net/).

3. Results and discussion

In the price series, presence of unit root can be expected, as evident from Fig. 1. To confirm suggestion incurred from visual inspection of the plotted data, presence of unit root was verified by KPSS test with auxiliary model including level constant. The unit-root test rejected the null hypothesis of stationarity in the price series at $\alpha = 0.05$. In the second step, a hypothesis of long-run cointegrating relationship was not proved by the Engle-Granger test. The outcome suggests that first difference transformation was required to satisfy the precondition of stationarity in the series. According to the procedure by Hyndman and Razbash (2014), single round of differencing was considered sufficient. Exploration of seasonality in the series showed that seasonal differencing was not necessary to remove potential seasonal unit root. To ascertain Granger causal relationships between the wheat prices, VAR(1) model was estimated in first differences.

Table 1. Coefficients of VAR(1) model estimated in first differences for time periods before EU accession and after.

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Before EU entry, T = 128</th>
<th>After EU entry, T = 112</th>
</tr>
</thead>
<tbody>
<tr>
<td>$dY_{bc}$</td>
<td><strong>0.326</strong></td>
<td><strong>0.345</strong></td>
</tr>
<tr>
<td>$dY_{af}$</td>
<td><strong>0.345</strong></td>
<td><strong>0.568</strong></td>
</tr>
<tr>
<td>$dY_{wm}$</td>
<td>0.034</td>
<td>0.097</td>
</tr>
<tr>
<td>$dY_{bc(t-1)}$</td>
<td>0.373</td>
<td><strong>0.316</strong></td>
</tr>
<tr>
<td>$dY_{af(t-1)}$</td>
<td><strong>0.230</strong></td>
<td><strong>0.210</strong></td>
</tr>
<tr>
<td>$S_1$</td>
<td>34.800</td>
<td>2.930</td>
</tr>
<tr>
<td>$S_2$</td>
<td>36.200</td>
<td>34.100</td>
</tr>
<tr>
<td>$S_3$</td>
<td>3.910</td>
<td>21.700</td>
</tr>
<tr>
<td>$S_4$</td>
<td>36.200</td>
<td>13.900</td>
</tr>
<tr>
<td>$S_5$</td>
<td>39.710</td>
<td>14.600</td>
</tr>
<tr>
<td>$S_6$</td>
<td>31.200</td>
<td>14.400</td>
</tr>
<tr>
<td>$S_7$</td>
<td>3.910</td>
<td><strong>0.316</strong></td>
</tr>
<tr>
<td>$S_8$</td>
<td>12.400</td>
<td><strong>214.600</strong></td>
</tr>
<tr>
<td>$S_9$</td>
<td>12.400</td>
<td><strong>214.600</strong></td>
</tr>
<tr>
<td>$S_{10}$</td>
<td>12.400</td>
<td><strong>214.600</strong></td>
</tr>
<tr>
<td>$S_{11}$</td>
<td>12.400</td>
<td><strong>214.600</strong></td>
</tr>
<tr>
<td>$S_{12}$</td>
<td>12.400</td>
<td><strong>214.600</strong></td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.486</td>
<td>0.537</td>
</tr>
<tr>
<td>DW</td>
<td>2.024</td>
<td>2.008</td>
</tr>
</tbody>
</table>

Notation: * significant at $\alpha = 0.1$; ** significant at $\alpha = 0.05$

Exploration of periodogram and output of the Fisher’s g-test indicated that periodic oscillations with 12-months frequency existed in the transformed series. It provided justification to incorporate additive seasonal constants in the VAR(1) model via binary indicator variables. Seasonal parameter $S_{12}$ was set to zero (Lütkepohl, 2005). Order of the VAR($p$) model was determined by likelihood ratio $\chi^2$-test and Schwarz’s information criterion (SIC), which tends to shield the VAR model from overfitting. In this study, the mentioned methods pointed to identical suitable lag of the VAR model $p = 1$ in the time segments before EU accession and after. VAR(1) model can be described by equation

$$y_t = S_t + \Phi_1 y_{t-1} + u_t.$$  

(2)
In this model, \( y_t \) denotes random vector of data observations, \( S_t \) denotes vector of fixed seasonal constants for the \( i \)-th period, \( \Phi_i \) is a \( k \times k \) square matrix of fixed mode coefficients, \( y_{t-1} \) denotes vector of lagged data observations and \( u_t \) represents \( k \)-dimensional innovation term. \( u_t \) follows white noise with nonsingular covariance matrix \( \Sigma_u \). We anticipated that short-term relationship among the named series in lagged form follows an unequal pattern in the period before Czech accession to EU and after. For this reason, specific VAR(1) models were estimated for the period before May 2004 and after. Parameter estimates of the vector autoregressive models are presented in Tab. 1.

In the VAR model, we found indication that price of wheat of human consumption quality (\( dY_{hc} \)) Granger-causes the price of wheat destined for animal feed (\( dY_{af} \)) in the period from September 1993 to April 2004. The respective parameter could be interpreted as expected change of CZK 0.345 in \( dY_{af} \) in the current month corresponding to one koruna change of \( dY_{hc} \) in the previous month. In terms of elasticity, the effect amounts to 0.397. In this model, significant 1-st order autoregressive effect of 0.326 was also observed in wheat price for human food quality (\( dY_{hc} \)). In terms of percent-to-percent change, it amounts to 0.332. Lagged \( dY_{af} \) also had small positive causal effect on \( dY_{hc} \). Autoregressive effect was also observed in \( dY_{wm} \), where it amounts to 0.316. Expressed in form of elasticity, it takes value of 0.285. Due to small size of the Czech market and restrictions in place for trade with agricultural commodities before 2004, lagged effects of \( dY_{hc} \) and \( dY_{af} \) upon \( dY_{wm} \) were deemed to have a casual nature. It could be concluded, that before EU entry, movements in lagged \( dY_{wm} \) had no significant impact upon \( dY_{hc} \) or \( dY_{af} \). Additionally, prices of wheat for human consumption and animal feed were determined by supply and demand in the closed Czech market and they influenced each other through a feedback system during this period.

In the period from May 2004 to August 2013, statistical evidence appeared that change in \( dY_{af} \) causes positive change the price of wheat for human consumption (\( dY_{hc} \)). The regression parameter could be interpreted as expected change of CZK 0.731 in \( dY_{hc} \) in the current month, corresponding to change in \( dY_{af} \) by one koruna, in the previous month. In terms of elasticity, the lagged effect amounts to 0.563. In this period, unit change in lagged \( dY_{wm} \) yields a positive effect of 0.210 upon \( dY_{hc} \). In terms of percent-to-percent change, the effect amounts to 0.253. In the VAR model, significant 1-st order autoregressive effect of 0.414 was found in wheat price for animal food quality. Corresponding percent-to-percent change amounts to 0.340. In addition, change in lagged \( dY_{wm} \) affects positively \( dY_{af} \) with the coefficient 0.193. It amounts to 0.259, in terms of elasticity. Shifted prices from the domestic market in the VAR(1) model showed no influence upon the current price in the world market. \( dY_{wm} \) price series appears to be 1-st order integrated with only weakly seasonally autocorrelated innovations.

To investigate size, significance and direction of response in one series to a positive innovation shock in another series, an orthogonal impulse response function (IRF) was calculated and plotted for prediction horizon 1 to 15 months. The magnitude of the innovation impulse in the source series was assumed one standard deviation. The orthogonal IRFs were computed according to Zivot and Wang (2006) and were accompanied by resampling 90% confidence intervals from 1000 bootstrapped replications. In a stationary VAR model, the IRFs are expected to converge to zero with time. In the second data segment after EU accession, shock in \( dY_{af} \) caused increase in \( dY_{hc} \) by approx. CZK 130 in the next month, followed by exponential decline to zero over the following seven months. A

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Fig. 2. (a) Response of \( dY_{hc} \) to innovation shock in \( dY_{af} \); (b) Response of \( dY_{hc} \) to innovation shock in \( dY_{wm} \); (c) Response of \( dY_{af} \) to innovation shock in \( dY_{wm} \).
positive impulse in $dY_{wm}$ caused a significant increase in $dY_{hc}$, peaking after two months. It later exponentially dropped to zero over the subsequent seven months. A similar response in $dY_{af}$ was observed to a positive shock in $dY_{wm}$. Diagrams of selected IRFs are provided in Fig. 2.

Graphical illustration of forecast error variance decomposition (FEVD) makes information available about relative impact of lagged variables in VAR(1) model upon prediction error variance in the response. Fig. 3 shows output of FEVD in $dY_{hc}$ and $dY_{af}$ series, since impact in the Czech domestic prices was mainly of our interest. The decomposition was obtained for a prediction horizon of 15 months. We noticed that it is primarily $dY_{af}$ series that determines the forecast error in predicted $dY_{hc}$ in horizon of 3 months or further (62%). The remaining share is accounted for by $dY_{wm}$ (22%) and $dY_{hc}$ (16%). In $dY_{af}$ series, on the other hand, the forecast error is 80% defined by movements in lagged $dY_{af}$ and only 20% by variation in $dY_{wm}$. Effects of lagged $dY_{wm}$ are too small and can be neglected. The breakdown of the forecast error variance relates to the VAR(1) model estimated from data segment subsequent to the Czech accession to EU.

![Fig. 3. (a) Forecast error variance decomposition of $dY_{hc}$; (b) Forecast error variance decomposition of $dY_{af}$.](image)

Assessment of estimated seasonal parameters indicates that a significant decline in wheat price happens every year shortly before new harvest comes to the market. It is followed by a price increase in the following two-to-three month period, after grain from new harvest becomes available. This seasonal pattern can be explained by necessity to vacate grain storage space in July and August before arrival of the new harvest. Fresh production is more valuable and therefore it is sold at higher price. In $dY_{wm}$ series, a different pattern of seasonality was detected with shortened 6-month frequency. It could be explained by a new harvest occurring twice a year in various parts of the world and thus affecting form of seasonality in the world market price.

The equations of differenced $dY_{hc}$ and $dY_{af}$ in VAR(1) model explained approximately 50% variability in the response variable and showed absence from serial correlation in the residual term. We noticed a better model fit ($R^2$) in regression equations for $dY_{hc}$ and $dY_{af}$ from the time segment following the Czech accession to EU. Identical relationship between lagged world wheat price and Czech wheat price was witnessed by Šyrovátku (2010), who also reported a similar pattern of seasonality in log-differenced monthly price series of prime quality wheat destined for human consumption.

4. Conclusions

Prior to EU entry, the Czech market with agricultural commodities was relatively isolated. Exports to other countries were possible, but licenses had to be obtained and volumes traded were limited. In this period, we observed no statistical evidence of world market price causally affecting fluctuations in prices in the Czech domestic market. Variation in Czech prices of wheat destined for human and animal consumption were determined by supply and demand. Within the sequestered national market, the prices affected each other in a feedback system. Both domestic prices also showed recurring seasonal movements throughout the calendar year. Form of seasonality was
tangled to wheat from the new harvest entering the commodity market. The average world market price could be described as 1-st order integrated series of innovations with short memory and seasonal components.

At present, Czech wheat producers take full benefits offered by free trade with cereals and other commodities within EU agricultural market. In this study, we found quantitative proof, that after EU accession, variation in monthly time series of two Czech wheat prices was explained by movements of average price in the world market lagged by one month. Since 2004, the level of Czech domestic prices of wheat progressively converges to the price level in the world market, as a consequence of dynamic trade between the Czech Republic and surrounding EU member states. It can be hypothesized, that convergence of prices will continue in the following years. As expected, the Czech domestic prices of wheat appear to have no important effect upon movements of price in the world market, primarily due to small size of the home market. The world market price could be characterized as 1-st order integrated series of innovations showing only weak seasonal autocorrelations. In the Czech market, the wheat price of animal feed quality causally affects the price of wheat destined for human consumption, but not vice versa.

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