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TESTING FOR MARKET INTEGRATION AND THE LAW OF ONE PRICE: AN APPLICATION TO SELECTED EUROPEAN MILK MARKETS

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ABSTRACT

This paper explores the long run linkages among milk prices of five European markets in a dynamic framework by employing multivariate cointegration analysis and appropriate Vector Error Correction (VECM) specifications. The detection of causal effects and the identification of possible dominant markets that drive the prices of other markets is carried out by means of Granger causality testing and exogeneity tests. Finally, the short run dynamics of the milk markets are explored by applying variance decomposition analysis.

JEL Classification: E320, F15, G150, Q110, C32, C51

Keywords: Milk; Law of one price (LOP); Market integration; EU; Cointegration

INTRODUCTION

The common policy for milk in the EU has its origins in 1964 (Council Regulation 804/68), and since then it has undergone several reforms. The EU dairy policy is based on a Common Organization of the Market (COM) whose main instrument is a target price for milk producers, supplemented by intervention instruments at both the supply and consumption of milk. These instruments include public intervention and private storage, internal subsidies for consumption and export refunds. The most important instrument is a "target price" and production control in the form of "milk quotas" first introduced in 1984[1]. During the Agenda 2000 negotiations four member countries, Denmark, Italy, Sweden and the United Kingdom, unsuccessfully campaigned for a dismantling of the quota system (Benjamin *et al.*, 1999).

The key elements in the current status of the common dairy policy regime after the last 2003 reform[2] are, first, the abolition of the Target Price, second, the introduction of a Single Payment Scheme (SPS) and thirdly, the facing out of quotas by 2015.

The Target Price for milk was originally introduced as a benchmark. Intervention prices were originally calculated from the Target Price at a level which would guarantee that milk prices for farmers would be maintained at a "reasonable level".

Since the use of intervention has declined, the Target Price eventually became redundant and it was abolished on 1 July 2004 as part of the reform[3].

The SPS is in the spirit of the general reform of the CAP: decoupled direct payments to farmers. Milk producers qualify for support payments to be paid per calendar year, per holding. The payments consist initially of two elements: dairy premiums paid equally to all milk producers; additional payments paid to milk producers according to criteria decided upon by the Member States.

Member states have the flexibility to introduce the SPS in 2005, 2006 or 2007. Dairy payments may be included in the SPS beginning in any one of these years. The SPS, including for the dairy sector, must be implemented by 2007. There is no option to continue with a combination of SPS and partially coupled payments from 2007 for dairy as there is for some other sectors[4].

Has the CAP, and the measures in its framework implemented hitherto, created fairly similar European prices for milk throughout the EU area? Are milk prices in different locations in Europe moving together? Is there at least a clear trend towards a unified price for milk? In other words, does the Law of One Price (LOP) prevail in the European milk markets? In this paper, we attempt to tackle these questions by testing for the LOP for milk prices in the EU.

Market integration and the LOP have been studied for a number of commodities. Richardson (1978), Ravallion (1986), Ardeni (1989) and Baffes (1991), tested the LOP for wheat, wool, beef, sugar, tea, tin, and zinc; Goodwin *et al.* (1990b), tested the LOP for a number of oilseed products, wheat varieties, corn, and sorghum; Sexton et al. (1991), tested for celery; Bellego (1992) and Sanjuan and Gil (1999) tested the LOP in the pork sector; Kadyrkanova *et al.* (2000), examined milk price linkages in Kyrgystan; Sanjuan and Gil (2001), tested the LOP in the European pork and lamb markets; Goodwin and Schroeder (1991a) and Goodwin (1992a, 1992b) evaluated the LOP in the wheat markets; Goodwin and Schroeder (1991b), tested for cattle; Zanias (1993), tested the LOP for four European Community products, namely, soft wheat, cow's milk, potatoes and pig carcasses; Gordon *et al.* (1993), evaluated the LOP in the European Community lamb market; Goletti and Babu, (1994) and Lutz *et al.* (1995), tested the LOP for maize; Silvapulle and Jayasuriya, (1994) and Ismet *et al.* (1998), tested for rice; Diakosavvas (1994), tested for beef; Jordan and VanSickle (1995), for fresh tomatoes; Froot et al. (1995), tested for grains (wheat, oats, and barley), dairy products (butter and cheese), eggs, peas, and silver; Kuiper et al. (1999), tested the LOP for corn.

So far, integration in the EU agricultural markets has been investigated through different methodological approaches. Firstly, by estimating static regressions among the involved price series (Tangeman, 1992). Secondly, by applying dynamic approaches based either on Granger causality testing (Blank, 1987; Gordon et al., 1993) or alternative causality procedures (Bellego, 1992). Thirdly, by applying Ravallion's (1986) approach (Dahlgran and Blank, 1992; Jordan and Vansickle, 1995). Fourthly, by means of Vector Autoregressive models (VAR) (Schroeder and Goodwin, 1990;

Goodwin and Schroeder, 1991a, and others). Finally, by means of cointegration techniques, using either a bi-variate cointegration framework (Ardeni, 1989; Baffes, 1991; Goodwin and Schroeder, 1991b; Zanias 1993) or a multivariate cointegration framework (Goodwin, 1992b; Silvapulle and Jayasuriya, 1994; Ismet et al., 1998; Sanjuan and Gil, 1999, and others).

In this paper, we explore the long run linkages among milk prices of five European markets in a dynamic framework by employing multivariate cointegration analysis and appropriate Vector Error Correction (VECM) specifications. The detection of causal effects and the identification of one or more possible dominant markets that drive the prices of the other markets are carried out by means of Granger causality testing and exogeneity tests. Finally, the short run dynamics of the milk markets is explored by applying variance decomposition analysis. The paper is organized as follows: Section 2 outlines the basic econometric tools applied in the empirical analysis. In Section 3, the results of the empirical analysis are reported and discussed. Finally the main conclusions are summed up in Section 4.

THEORETICAL AND METHODOLOGICAL ISSUES

Integration Analysis

It is well known that in the case of non-stationary data series the results of an econometric analysis series are spurious because the classical **t** and **F** tests proved improper (Fuller, 1976). Consequently, the first step is to test the series for stationarity and determine the order of integration of the examined variables. In this context, all the employed variables were tested for unit roots utilizing the Augmented Dickey-Fuller (ADF) test. Using the Ordinary Least Squares (OLS) procedure the ADF test estimates the following equation:

$$\Delta y_t = \delta y_{t-1} + \Sigma\, \delta_i \Delta y_{t-i} + \varepsilon_t \qquad \sum_{i=1}^{k} \tag{1}$$

Where $\Delta y_t = y_t - y_{t-1}$ and t = time trend variable

Hereafter the hypothesis $\delta = 0$ against the alternative $\delta < 0$ was tested. Rejection of the null hypothesis implies that $y_t \sim I(0)$.

Cointegration

The long-run relationship between a number of series can be looked at from the viewpoint of cointegration (Engle and Granger, 1987). Cointegration is a time series modeling technique developed to deal with non stationary time series in a way that does not waste the valuable long-run information contained in the data. Moreover, the need to evaluate models which combine both short-run and long-run properties and which, at the same time, maintain stationarity in all of the variables, has prompted a reconsideration of the problem of regression using variables measured in their levels.

As Granger and Newbold (1974), and Phillips(1986 and 1987) pointed out, given that many economic time series exhibit the characteristics of integrated processes of order one, I(1), estimating traditional OLS or VAR models with I(1) processes can lead to nonsensical or spurious results. Note that, I(1) processes are those which need to be differenced to achieve stationarity.

Let x(t) be a vector of n-component time series each integrated of order one. Then x(t) is said to be cointegrated CI(1, 0), if there exists a vector ϕ such that

$$s(t) = \phi' x(t) \tag{2}$$

is I(0). Stationarity of s(t) implies that the n variables of x(t) do not drift away from one another over the long-run, obeying thus an equilibrium relationship. If ϕ exists, it will not be unique, unless x(t) has only two elements. The Engle and Granger approach can deal with the possibility of only one linear combination of variables that is stationary. Recent advances in cointegration theory (Johansen and Juselius, 1990) have developed a maximum likelihood (ML) testing procedure on the number of cointegrating vectors which also allows inferences on parameter restrictions. The ML method uses a vector autoregressive (VAR) model

$$\Delta x(t) = \sum_{i=1}^{q-1} \Pi_i \Delta x(t-i) + \Pi_q x(t-q) + \mu + v(t) \tag{3}$$

where x(t) is a n×1 vector of variables, Πq is a n´n matrix of rank r≤n, μ is a n×1 vector of constant terms, v(t) is a n×1 vector of residuals and Δ is the first difference operator. The testing procedure involves the hypothesis H_2: $\alpha\beta'$, where α and β are n×r matrices of loadings and eigenvectors respectively, that there are r cointegrating vectors β_1, $\beta_2, \ldots, \beta_r$ which provide r stationary linear combinations β'x(t-q). The likelihood ratio (LR) statistic for testing the above hypothesis is given below:

$$-2InQ = -T \sum_{i=r+1}^{n} In(1 - \hat{\lambda}_i) \tag{4}$$

This ratio is a test that there are at most r cointegrating vectors versus the general alternative (trace), where λ_i corresponds to the n-r smaller eigenvalues. The n×r matrix of cointegrating vectors β can be obtained as the r, n-element eigenvectors corresponding to λ_i. The LR test statistic for testing r against r+1 cointegrating vectors is given by

$$-2In(Q : r \mid r+1) = -T \cdot In(1 - \hat{\lambda}_{r+1}) \tag{5}$$

The above tests (4) and (5) are used to determine the significant eigenvalues and the corresponding number of eigenvectors.

Generalized Forecast Error Variance Decomposition

The forecast error variance decomposition provides a decomposition of the variance of the forecast errors of the variables in a VAR model at different time horizons. In this paper, we use the generalized forecast error variance decomposition which for the i-th variable in the VAR is given by Pesaran and Pesaran, (1997).

$$\psi_{ij,N} = \frac{\sigma_{ii}^{-1} \sum_{k=0}^{N} (e_j' A_k \Sigma e_i)^2}{\sum_{k=0}^{N} e_i' A_k \Sigma A_k' e_i}, \tag{6}$$

where Σ is the covariance matrix of the shocks $\mathbf{u}_t$ in the considered VAR; $\mathbf{e}_i$ is the selection vector defined by $\mathbf{e}_i$=(0,0,...0,1,0,...0)′ with 1 the i-th element; and $\mathbf{A}_k$, k=0,1,2,... are the coefficient matrices in the moving-average representation of the VAR model. In (6), $\psi_{ij,N}$ measures the proportion of the variance of the N-step forecast errors which is explained by conditioning on the non-orthogonalized shocks, u_{it}, $u_{i,t+1}$, ..., $u_{i,t+N}$, but explicitly to allow for the contemporaneous correlations between these shocks and the shocks to the other equations in the system.

DATA AND EMPIRICAL RESULTS

Data

The LOP is tested for five European milk markets, namely France, Germany, Denmark, Netherlands and Belgium. The selection of the above countries was based on data availability. More particularly, the data for the respective five price series, denoted by LFR, LGER, LDK, LNTH and LBG are monthly, refer to selling prices of raw cows' milk, 3.7% fat content and cover the period from January 1980 to December 2003. All prices are Euro-fixed per 100 Kg and are used in logarithmic form. Data source is Cronos Prague databank.

Integration and Cointegration Analysis

The results reported in Table 1, indicate that the null hypothesis for the existence of a unit root could not be rejected at 5% significance level and thus none of the series is stationary when the test refers to the log-levels of the variables. Next, the above tests were applied on the first differences of the logarithms of the series. The results, reported in Table 1 as well, provide evidence that all the examined variables are integrated of order one I(1).

Given that all the respective variables were found integrated of order one I(1), we proceeded with the Johansen (1988) multivariate cointegration procedure in order to detect possible long run relationships among the considered price series. It should be noted that the number of cointegrating vectors indicates the degree of market integration. In this sense, the more cointegrated price pairs are found the stronger the degree of integration among the national markets for milk.

Table 1
Augmented Dickey-Fuller (ADF) Unit-Root Tests

Variable	*Levels*	*First Differences*
LGER	-2.191 (6)	-10.394 (6)
LFR	-2.766 (6)	-17.099 (6)
LDK	-3.386 (3)	-8.019 (3)
LNTH	-3.066 (6)	-12.698 (4)
LBG	-3.015 (5)	-10.827 (3)

Note: The ADF statistics were calculated with a number of lags (indicated in the parentheses) to ensure that the residuals were "white noise". The critical value from Fuller, for the respective degrees of freedom and the 5% level of significance is -3.4272.

The results from both the "maximal eigenvalue of the stochastic matrix" test as well as the "trace" test are presented in Table 2. The reported results reveal the existence of four cointegrating vectors and hence a single common trend which leads the set of price series. This finding suggests that the examined EU milk markets are strongly interdependent and the degree of market integration may be considered

Table 2
Cointegration Results

281 observations from 1980M1 to 2003M5.
List of variables included in the cointegrating vector:LFR LGER LDK LNTH LBG
List of I(0) variables included in the VAR: S1 S2 S3 S4 S5 S6 S7 S8 S9 S10 S11

Panel A: Cointegration LR test based on maximal eigenvalue of the stochastic matrix

Null hypothesis	*Alternative hypothesis*	*Test statistic*	*Critical value(95%)*	*Critical value (90%)*
$r = 0$	$r = 1$	48.4701	29.9500	27.5700
$r \leq 1$	$r = 2$	39.7241	23.9200	21.5800
$r \leq 2$	$r = 3$	21.0319	17.6800	15.5700
$r \leq 3$	$r = 4$	13.5682	11.0300	9.2800
$r \leq 4$	$r = 5$	4.1177	4.1600	3.0400

Panel B: Cointegration LR test based on trace of the stochastic matrix

Null hypothesis	*Alternative hypothesis*	*Test statistic*	*Critical value (95%)*	*Critical value (90%)*
$r = 0$	$r \geq 1$	126.9120	59.3300	55.4200
$r \leq 1$	$r \geq 2$	78.4419	39.8100	36.6900
$r \leq 2$	$r \geq 3$	38.7178	24.0500	21.4600
$r \leq 3$	$r \geq 4$	17.6859	12.3600	10.2500
$r \leq 4$	$r = 5$	4.1177	4.1600	3.0400

Note: Critical values are taken from Johansen and Juselius (1990).

"perfect". The estimates of the long-run relationships among the considered milk price series are reported in Table 3.

Table 3
Estimated Cointegrated Vectors (Normalized in Brackets)

	Vector 1	*Vector 2*	*Vector 3*	*Vector 4*
LFR	.61266	-.57252	-.14318	-.13902
	(-1.0000)	(-1.0000)	(-1.0000)	(-1.0000)
LGER	-.83249	-.042333	-1.2742	-1.4910
	(1.3588)	(-.073942)	(-8.8991)	(-10.7252)
LDK	-.68800	1.7786	1.1080	.41353
	(1.1230)	(3.1067)	(7.7387)	(2.9747)
LNTH	.57454	-1.8122	.027755	2.1763
	(-.93778)	(-3.1653)	(.19384)	(15.6550)
LBG	.67475	-.49825	-.49825	-1.2799
	(-1.1013)	(-.87027)	(-2.0047)	(-9.2065)

Long Run Exogeneity

In order to explore if there is some price leading or driving the joint evolution of the system in the long run, we investigated each price for weak exogeneity by carrying out tests on the significance of the adjustment coefficients, that is, the elements in matrix **a**. For example, the null hypothesis of weak exogeneity for the first considered variable, the milk price in Germany (LGER), can be formulated as follows:

H_0: weak exogeneity of price LGER:

or

$$H_0 : \mathbf{a} = \begin{array}{c} \begin{bmatrix} 0 & 0 & 0 & 0 \\ * & * & * & * \\ \vdots & \vdots & \vdots & \vdots \\ * & * & * & * \end{bmatrix} \end{array} \begin{array}{c} 1 \\ 2 \\ \vdots \\ r \end{array}$$

With r to denote the number of the considered endogenous variables.

The results, presented in Table 4, indicate that in all cases the null of weak exogeneity is rejected. The latter implies that none of the investigated milk markets seem to have more autonomy in the pricing determination process or it plays a dominant role in price formation in the long run.

Table 4
Likelihood Ratio Tests of Long run Weak Exogeneity

Price variable	*LGER*	*LDK*	*LNTH*	*LFR*	*LBG*
Likelihood Ratio Statistic	14.5244	15.9162	19.6354	26.7625	24.0704
p-value	0.005	0.003	0.000	0.000	0.000

SHORT RUN DYNAMICS

Granger Causality Analysis

With regard to the detection and the direction of possible causal effects among the examined series we proceeded with testing the implied error-correction models (ECM) for Granger-causality and next by applying innovation accounting analysis.

The results reported in Table 5 reveal the followings: In the short run, both France and Denmark, Granger-cause milk prices in Germany and Belgium. Netherlands is causally affected from both Germany and Belgium while feedback effects are detected between Germany and Belgium.

Table 5
Granger-Causality Tests

	Hypothesis tested		*Wald statistic*	*p-value*
LFR	does not cause	LGER	37.113	0.000
LNTH	>>	LGER	11.294	0.586
LDK	>>	LGER	22.738	0.045
LBG	>>	LGER	23.059	0.027
LGER	does not cause	LFR	15.209	0.230
LNTH	>>	LFR	6.982	0.859
LDK	>>	LFR	12.678	0.393
LBG	>>	LFR	12.585	0.400
LGER	does not cause	LNTH	39.782	0.000
LFR	>>	LNTH	16.168	0.240
LDK	>>	LNTH	11.139	0.599
LBG	>>	LNTH	23.578	0.035
LGER	does not cause	LDK	18.870	0.127
LFR	>>	LDK	19.419	0.111
LNTH	>>	LDK	17.578	0.174
LBG	>>	LDK	18.051	0.156
LGER	does not cause	LBG	37.379	0.000
LFR	>>	LBG	52.199	0.000
LNTH	>>	LBG	15.645	0.269
LDK	>>	LBG	25.127	0.022

Variance Decompositions Analysis

Sims (1980), proposed the use of variance decomposition analysis mainly for economic policy evaluation since it provides a means for out of sample forecasting the effects of a shock. In this section, the variance decomposition method is applied and results are derived. The moving average representation can be used to depict the responses of all variables to shocks (i.e. innovations) in the residuals. Given the unrestricted VAR system, typical random shocks are positive residuals of one standard deviation unit in each equation. The relevant analysis describes the effect of a one standard deviation shock to the residuals.

Following Pesaran and Pesaran (1997), the *generalized* forecast error variance decompositions of the milk price series for all the considered countries were calculated and are reported in Table 6. More specifically, this table reports the percentage of the variance of the k-month ahead forecast error of the variables that is attributable to each of the shocks for k=3, 6, 12, 18, and 24. We consider a 6-months ahead time horizon as short-run, a 12-months ahead time horizon as medium-run and a 24-months ahead horizon as long-run.

Table 6
Forecast Error Variance Decompositions

Panel A: Generalized Forecast Error Variance Decomposition for Variable *LFR*

Forecast horizon	*Percentage of variance of error due to innovations in* *LFR*	*LGER*	*LDK*	*LNTH*	*LBG*
3	.90597	.15664	.09088	.008951	.011188
6	.82701	.18802	.16601	.018107	.030601
12	.76238	.21360	.22271	.018778	.055544
18	.69289	.25587	.26893	.024956	.035682
24	.64525	.25102	.32025	.021530	.029424

Panel B: Generalized Forecast Error Variance Decomposition for Variable *LGER*

3	.047904	.91211	.07407	.10603	.21656
6	.058884	.88520	.13308	.10899	.22607
12	.082060	.80483	.27362	.10525	.18161
18	.073917	.73503	.37269	.10491	.15266
24	.071476	.69363	.40659	.08890	.12495

Panel C: Generalized Forecast Error Variance Decomposition for Variable LDK

3	.052666	.20549	.91356	.006010	.029905
6	.057603	.28225	.85253	.037141	.037862
12	.052911	.36783	.75063	.096421	.038799
18	.053758	.42843	.70593	.103100	.044348
24	050708	.44698	.69142	.098890	.035604

Panel D: Generalized Forecast Error Variance Decomposition for Variable *LNTH*

3	.017236	.08684	.02976	.94052	.12306
6	.024434	.18708	.06207	.78341	.21429
12	.018295	.32217	.15515	.61344	.17507
18	.015683	.30062	.27868	.54546	.14509
24	.013053	.27942	.31497	.45633	.14799

Panel E: Generalized Forecast Error Variance Decomposition for Variable LBG

3	.044533	.20802	.12109	.05138	.87330
6	.067623	.33772	.18737	.07688	.73342
12	.058138	.31161	.29178	.11957	.63809
18	.047594	.29693	.38248	.14297	.55421
24	.050533	.27347	.42644	.14425	.48943

In the case of France, we observe that only Germany and Denmark efficiently explain the milk price variance in all time horizons with a percentage between 20 and 25 for each one of them. For Germany, only Denmark seems to have explanatory power which turns significant (over 30%), after the 12-month horizon. With regard to Denmark, Germany is the only explanatory factor over the whole time horizon, though the most significant effects reveal after the 12^{th} month (30-40%). Finally, for both the Netherlands and Belgium, Germany and Denmark are found, after the 12-month horizon, to exert significant impacts, explaining each one, about 25%-30% of the price variance.

CONCLUDING REMARKS

In this paper, we explore the long run linkages among milk prices for five European markets in a dynamic framework by employing multivariate cointegration analysis and appropriate Vector Error Correction (VECM) specifications. The detection of causal effects and the identification of one or more possible dominant markets that drive the prices of the other markets were carried out by means of Granger causality testing and exogeneity tests. Finally, the short run dynamics of the milk markets was explored by applying variance decomposition analysis.

The results reveal the existence of a single common trend which leads the set of price series. This finding suggests that the examined EU milk markets are strongly interdependent and the degree of market integration may be considered "perfect". In other words, common European agricultural policy achieved the unification of milk markets and the adoption of similar prices throughout the EU area. With regard to the short run dynamics of the involved milk price series, there is evidence that German and Denmark milk markets dominate in Europe and drive the milk prices of the other markets.

NOTES

1. Council Regulation (EEC) No 856/84, and Council Regulation (EEC) No 857/84, were replaced by Regulation (EEC) No 3950/92, which was later amended by Regulation (EC) No 1256/1999 in the framework of Agenda 2000.
2. CEC Regulations 1782/2003; 1787/2003; and 1788/2003.
3. The abolition of the Target Price required a further change, concerning the superlevy - the penalty when a country exceeds its quota. This was used to be expressed as 115 % of the target price. As the target price has been abolished, the new dairy regulation states the specific value of the superlevy as: - EUR 33.27/100 kg for 2004/05; - EUR 30.91/100 kg for 2005/06; - EUR 28.54/100 kg for 2006/07; - EUR 27.83/100 kg for 2007/08 and subsequent periods.
4. The total amounts available for direct dairy premiums in a given year are based on quota held at the end of the preceding quota year and are as follows: EUR 8.15; 16.31; 24.49/tonne of quota for calendar year 2004; 2005; and 2006 respectively.

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