Redefining integration of the European dairy markets: do they converge?

Christos Karelakis

Democritus University of Thrace, Department of Agricultural Development, Pantazidou 193, 68200, Orestiada

** Corresponding author: e-mail:* chkarel@agro.duth.gr

Abstract

The European dairy market has received inefficient protection through the years within the context of significant policy reforms. Until the next reform, the various protection measures will be gradually dismantled, fostering the competitiveness of the sector. The present paper exploits the prospects of an integrated European dairy market with the investigation of the price behaviour for EU countries through the Purchasing Power Parity (PPP) and the assistance of the Johansen cointegration technique. Three major groups of dairy markets have been identified for which a single cointegration vector is confirmed. The results validate the existence of a long-run relationship among real exchange rates based on dairy prices for these groups of countries, whereas, weakly exogenous variables existed across all groups, offering the other countries within each group, protection against external shocks and adversaries of price volatility.

Keywords: cointegration, Purchasing Power Parity, dairy markets, convergence

1. Introduction

In the current turbulent economic environment, the dairy sector of the European Union (EU) follows a transitional path from high protection regulation to international and domestic liberalisation. The Common Agricultural Policy (CAP) has inefficiently protected the sector (and apparently national and regional markets) through a complicated system of price support, production quotas, import restrictions, and export subsidies. These policy schemes are being dismantled, with the milk delivery quota system, in particular, coming to an end in the year 2015. The abolition of production restrictions has been widely welcomed since it may foster Europe's relative global competitiveness. Recent data indicate that the sector produces almost 36% of the world's dairy production, conducting 27% and 14% of the world's dairy exports and imports respectively. Its economic significance is also evident from the substantial contribution (13%) to the agricultural turnover in the EU, providing around 10% in the Food and Drinks industry, while it employs approximately 10% of total employment based on 24 million cows in approximately one million farms (IUF 2012).

Disassembling the protection measures will possibly increase dairy production, raising more concerns about oversupply and possible weakened commodity prices. Prices have already dropped in the EU after reaching a peak in 2007, affecting dairy producers' income. It seems likely that the EU dairy prices will closely align with international prices, which, however, are volatile and this may be transmitted to EU prices. The increased volatility, along with the weakened prices and the absence of subsidies, will guide farmers to viability risk. The EU Commission acknowledges these trends, confronting the opportunity to implement policy schemes that will ensure the long-run competitiveness of the sector (IUF, 2012). Notably, through a transitional approach of further liberalising the intra EU trade that will expedite dairy market integration, allowing dairy farmers to be more market-oriented and consumers to benefit from safe and quality dairy products.

An important condition for a perfectly integrated market is the validity of the law of one price (LOOP) or Purchasing Power Parity (PPP) for a market, which has been empirically confirmed only in the short-run (Goldberg and Verboven, 2001). The backbone of the CAP policy framework prompts to attain market integration for agricultural products, for simultaneously determine prices in all countries and furthermore, amendments of one price for various reasons to be transmitted to other countries. Nevertheless, the effectiveness of a specific policy scheme is controlled by different market characteristics, including the extent of market power, the existence or absence of barriers to entry, the product differentiation etc.

Drawing attention to the considerations mentioned above, the present study examines the existence of an integrated European dairy market given that the dairy products sector is one of the most important agricultural sectors in the EU. The objective is to get an insight into real exchange rates based on the prices for dairy products, to indirectly investigate the EU dairy market integration. This is achieved through the investigation of the price behaviour for EU countries through the Purchasing Power Parity (PPP) and the assistance of the Johansen cointegration technique. Thus, an endeavour is made to examine whether the implementation of the reformed CAP policy schemes promote a standard behaviour for the domestic dairy prices against the dairy prices of a foreign country (in our case the USA was chosen).

The remainder of the paper begins with a brief description of the European dairy market within the CAP framework. Subsequently, the data and the methodology employed are presented, followed by the empirical results. Finally, the results and their implications are discussed within the spectrum of the existing policy context, and conclusions are drawn.

2. The CAP and the EU dairy market

The overarching objective of the EU economic policies is market integration, which also reflects the primary goal of the CAP through the Common Market Organizations (CMO's). The procedure towards integration developed utilizing removal of trade barriers, harmonization of tax rates and other national regulations, increased transparency, monitoring of cross-country price differences and reduction of exchange rate volatility. The spatial market integration can effectively offer a clear insight into the intricacies of the market mechanism and behaviour. A key element for a well-integrated market is to achieve a unique equilibrium for prices attributed to the competition for arbitrageurs, while the differences in prices are related to differences in transportation and transaction costs. Actually, according to Sexton et al., (1991), market integration serves as an indicator of competitiveness, the efficacy of arbitrage and price efficiency. On the other hand, weak or non-integrated markets may well send misleading signals, generating market imperfections, distortions and inefficiencies. Further, non-integrated markets lag in responding to price signals that bring about the failure to exploit opportunities in other markets.

The configuration of the CAP measures serves the purpose of a single integrated market for all agricultural products, where common prices will exist. Theoretically, based on economic theory, the law of one price (LOOP) should hold in such a market (O' Connell, 2002). Nevertheless, despite the significant reforms that the CAP has undergone and in conjunction with the recent EU enlargement, the existed barriers for the agricultural markets could not be eliminated, primarily attributable to various factors related to inefficient or even dearth of arbitrage mechanisms. For instance, imperfect information or risk aversion may well lead to inefficient arbitrage, while imperfect competition in these markets plays a defensive role in market integration (Cambell and Hopenhayn, 2005; Zanias, 1993).

The potential impact of the CAP reform on dairy markets is contingent upon the demand evolution for dairy products in the EU (Bouamra-Mechemache et al., 2008) and considering that dairy production is regulated by quota, the market price equilibrium is mainly affected by demand; it is relatively price inelastic though more sensitive to changes in income The Fischler reform of the CAP in 2003 substantially modified the dairy market's subsidy system and farmers confront an unprecedented situation with soaring prices for agricultural raw materials, selling their products at a higher price (dairy, meat and cereals), whilst coping with increased prices of concentrates. A further decrease in prices for dairy products is also expected, resulting in price levels below those projected in the baseline for all products except for skimmed milk powder (SMP). Lower internal prices mean that there will be fewer surpluses to export that deteriorate the terms of trade for subsidized exports.

Nevertheless, the positive notion of the recent CAP reform is to induce the adjustment of dairy production to market demand through the shift from price support mechanism to decoupled farm payments. In this way, the dairy sector is now set on a more market-oriented course within which farmers are invited to survive and prosper. Ultimately, the EU dairy process will adjust more to the world prices, consequently increasing both the price volatility and the farmer's risk to maintain income and the farm's economic viability. Theoretically, such policy measures stimulate the concept of profitable agricultural enterprises, while the level of subsidy payments cease to be the production determinant. The concept of real business is starting to be applied equally in the case of agribusinesses, even though the contribution of subsidy payments to farm income is not taken into consideration when evaluating market returns. This situation is leading to calls for market liberalization and integration, weakening the pricing support mechanism.

3. Materials and methods

3.1. The data

The present survey employs consumer price indices of dairy prices for the EU and the USA and the nominal exchange rates in order to calculate the real exchange rates. The calculation is based on the following formula:

$$q_t = s_t + p_t - p_t^* \qquad \textbf{(1)}$$

where q_t is the real exchange rate

s_t is the nominal exchange rate

p_t is the domestic consumer price of dairy products for the domestic country

p_t^* is the domestic consumer price of dairy products for the foreign country (USA)

The dataset was derived from Eurostat (the consumer price indices for dairy products of EU countries), the USDA (the consumer price indices for dairy products of USA) and the International Monetary Fund (nominal exchange rates - euros/dollar) covering the period from January 1996 to December 2011. Initially, we employed the unit root test to survey the degree of integration for each time series employed, and afterwards, with the assistance of Johansen cointegration technique, we examined whether the real exchange rates are co-integrated. The application of the Johansen cointegration technique preceded the estimation of a general linear mixed model on our data to assess the effects of TIME (monthly prices of dairy) and COUNTRY. Subsequently, we derived the residuals on which we employed the Johansen cointegration technique. The next step involved the estimation of the Vector Error Correction Model to capture the short-term dynamics for every identified group of countries, as well as the Granger causalities among the members of each group. The final step included an impulse response analysis to survey the reaction of the real exchange rates of the countries within the same group to an innovation leading to a change of one standard deviation to one of the countries of the particular group.

3.2. Methodology

3.2.1. The mixed models

Longitudinal (generally clustered data) typically arise in biometrical and epidemiological situations, where the experimental units measured repeatedly over time are humans. However, longitudinal data may occur in other research contexts (i.e., social and economic) where several statistical methods may apply for their analysis. The recent statistical literature pinpoints two modern methods: the random effects model (Laird and Ware, 1982), also known as the general linear mixed model (GLMM), and the generalized estimating equations (GEEs) (Liang and Zeger, 1986). Both methods comprise a part of the broader class of techniques called generalized linear models (GLMs).

In the current context, the nature of the data collected is longitudinal, in a sense that observations on dairy prices within each country are typically correlated, bearing also in mind the possible variations among the 22 European countries in our dataset. The possible correlation of the within-cluster (in our case within-country) measurements is an essential factor in the analysis of longitudinal data and the incorporation of the (possible) serial correlation of within-subject measurements into the model consists a key advantage.

A multilevel-type analysis was conducted to investigate and explore the complex ways in which time or country effects - as well as their interactions - can shift, beyond merely measuring changes in average aggregate European dairy prices. With only two sources of variation (Level 1: YEAR, Level 2: COUNTRIES) the fit of a General Linear Mixed Model allowing two sources of variation was chosen as a suitable modelling approach. Subsequently, longitudinal data analysis methodology was employed to assess the effects of time (yearly prices of dairy) and Country on dairy prices. We also looked at differences in terms of the interaction between the two variables. We focused our attention on finding evidence of significant effects of [COUNTRY] and [TIME] parameters on the outcomes.

National variations were of particular interest: ignoring either the within-country [TIME EFFECT] or the between-country variation in the data can affect the analysis of the data and, for example, may lead to the underestimation of the standard errors of regression coefficients. Furthermore, if the researcher measures a single source of variation, for instance, the only variation among European countries, he/she may lose significant information attributed to the between countries differences. In addition, since one may assume that the measurements taken within the same country are more likely to be correlated, compared to measurements taken from different countries, the within-country correlation in the data should also be taken into consideration. For this reason, the variable [COUNTRY] has been included in the fitted GLMM model as a possible covariate.

The General Linear Mixed-(Effects) Model for Longitudinal Data, initially proposed by Laird and Ware (1982), can be written as:

$$\mathbf{y}_i = \mathbf{X}_i\mathbf{b} + \mathbf{Z}_i\mathbf{u}_i + \boldsymbol{\varepsilon}_i\,,\ (i = 1,2,...,m)$$

where:

- $\mathbf{y}_i = \left(y_{i1}, y_{i2},..., y_{in_i}\right)^t$ denotes the $(n_i \times 1)$ vector of responses for the i_{th} subject
- $\mathbf{X}_i$ is a $(n_i \times p)$ design matrix that characterizes the systematic part of the response, e.g. depending on covariates and time.
- $\mathbf{b}$ is a $(p \times 1)$ vector of (population-specific) fixed parameters, namely the fixed effects.
- $\mathbf{Z}_i$ is a $(n_i \times q)$ design matrix that characterizes random variation in the response due to among-unit sources.
- $\mathbf{u}_i$ is a $(q \times 1)$ vector of (subject-specific) random effects and finally,
- $\boldsymbol{\varepsilon}_i$ is a $(n_i \times 1)$ vector of within-unit errors, usually called random error.

Since each subject's response vector $\mathbf{y}_i = \left(y_{i1}, y_{i2},..., y_{in_i}\right)^t$ consists of n_i repeated measurements, it is evident that the total number of observations included in the longitudinal model will be $N = \sum_{i=1}^{m} n_i$, in total. As concerns the distributional behaviour of the random terms of the mixed effects model, it is customary to specify a fully parametric form for both the subject-specific random effects $\mathbf{u}_i$ $(i = 1,2,...,m)$ and the random errors $\boldsymbol{\varepsilon}_i$ $(i = 1,2,...,m)$. Normality is the most common parametric assumption for the distribution of $\mathbf{u}_i$ and $\boldsymbol{\varepsilon}_i$ (Laird and Ware, 1982), that is we assume:

$$\mathbf{u}_i \sim N_q(\mathbf{0}, \mathbf{D}) \text{ and } \boldsymbol{\varepsilon}_i \sim N_{n_i}(0, \mathbf{R}_i),$$

in which $\mathbf{D}$ is a $(q \times q)$ variance-covariance matrix that characterizes variation due to between-subjects sources, and $\mathbf{R}_i$ a $(n_i \times n_i)$ variance-covariance matrix that characterizes variance and correlation due to within-subjects sources (i.e. the variation that occurs due to measurement error or due to biological within-unit fluctuations). Also, it is assumed that the $\mathbf{u}_i$'s and $\boldsymbol{\varepsilon}_i$'s are distributed independently for $i = 1,2,...,m$. Further, notice that the above model assumes homogeneity of variance only for the $\mathbf{u}_i$'s (constant variance-covariance matrix $\mathbf{D}$ for all subjects $i = 1,2,...,m$).

The most common choice for the (within-subject) variance-covariance matrix $\mathbf{R}_i$ in many applications is the simple covariance structure:

$$\mathbf{R}_i = \sigma^2 \mathbf{I}_{n_i},$$

where $\mathbf{I}_{n_i}$ denotes the $(n_i \times n_i)$ identity matrix and σ^2 is a (unique) variance parameter used to describe the within-subjects variability in the data. This parameterization suggests that the variance is the same across each ith $(i = 1,2,...,m)$ individual's separate measurements $y_{i1}, y_{i2},..., y_{in_i}$ and furthermore, these measurements were taken sufficiently far apart in time so that the possibility of correlation among them is practically considered negligible. The GLMM with this specific additional restriction is called the conditional-independence model.

3.2.2. Selection of the covariance structure

The feasibility to choose among a broad variety of covariance structures for the GLMM is one of the strategic advantages that this model accommodates in the analysis of longitudinal data. By selecting a structure that best fits the real data covariance results in obtaining the best possible efficient estimates of fixed effects. The mixed model analysis allows for substantial flexibility in specifying the correlation structure within cases and offers the potential for valuable substantive insights into the nature of that correlation. Each structure is suitable for describing the within-subjects covariance structure based on the specific dataset. For example, if the correlations at all lags are of about the same magnitude, then a symmetric compound structure seems reasonable to describe the within-subjects covariance. If the correlations are shown to decay exponentially with the time lag, then an autoregressive covariance structure can be considered as the most appropriate to describe the dependence of observations within a subject (Verbeke and Molenberghs, 2000).

We chose to fit models assuming four (4) different variance-covariance structures [the independent or diagonal, the compound symmetric, the Toeplitz and the autoregressive of order 1-AR(1)] to select the specific structure that best fits the possible within-subject correlation of our dairy prices data. Table 1 illustrates the values of the goodness of fit criteria[1] for selecting the best covariance structure from which it is clear that the best fit (smallest values) corresponds to the AR(1) covariance matrix, whereas the worst model corresponds to the one assuming independence among the within-country measurements.

[1] AIC: Akaike's Information Criterion, AICC: Hurvich and Tsai's Criterion, CAIC: Bozdogan's Criterion, BIC: Scharz's Bayesian Criterion

Table 1: *Goodness-of-fit values for selection of the best covariance structure*

	-2REML	AIC	AICC	CAIC	BIC
Diagonal	10289.042	10505.042	10515.446	11236.363	11128.363
Compound symmetric	310.258	314.258	314.263	327.801	325.801
AR(1)	-5071.445	-5067.445	-5067.440	-5053.902	-5055.902

* Toeplitz covariance matrix was not found applicable to the current analysis

3.2.3. Unit Root tests and the Johansen Cointegration technique

The methodology prerequisites the investigation of whether the time series are I(1), which means that they are non-stationary in levels and stationary in first differences. Stationarity was surveyed with the application of the widely used method for testing the existence of a unit root in the time series, the ADF unit root test (Dickey and Fuller, 1979). The objective is to test the null hypothesis for a single unit root in the data - generating a process for any variable surveyed. The Akaike and the Schwartz - Bayesian (SBC) criterion was employed to determine the ADF form, choosing the model with the lowest value in the specific criterion, for every time series. Subsequently, the final form of the auxiliary regression that includes was determined a constant and a time trend for all the variables employed, concluding that the time series examined are either I(1) or I(0) and so their combination can be tested for stationarity with the application of Johansen cointegration technique. In this case, the variables studied are cointegrated, and hence, there is a long-run relationship between them.

The power of the ADF test was in many cases in doubt and Elliot et al., (1992) suggested a simple modification, namely the DF-GLS test. The test is shown to be approximately uniformly most power invariant (UMPI), while no strictly UMPI test exists. The reported Monte Carlo results indicate that the power improvement from using the modified Dickey-Fuller test can be substantial, although the particular methodology derived the same results.

The Johansen cointegration technique involves testing the null hypothesis that there is no cointegration against the alternative that there is cointegration. The method employs two likelihood ratio (LR) test statistics (the trace and the maximal eigenvalue (A-max) statistics), to test for the number of cointegrating vectors in non-stationary time series. The critical values are taken from Osterwald and Lenum (1992), which differ slightly from those reported in Johansen and Juselius (1990). The application of the technique presupposes the calculation of the number of lags of the model's endogenous variables since an autoregressive coefficient is used in modelling each variable. The determination of the number of lags depended on the Akaike information criterion and the Ljung - Box test. Based on the Granger representation theorem, if a cointegrating relationship exists among a set of I(1) series, a dynamic error-correction (EC) representation of the data also exists and therefore, in a second stage we estimated the

Vector Error Correction Model to examine the direction of the causality between the four variables employed. This direction is determined by the statistical significance of the cointegrating equation coefficient. Additionally, the error correction model captures not only the long-term but also the short-term dynamics of the model.

4. Results and Discussion

The fixed effect of the GLMM models were estimated prior to the implementation of the Johansen cointegration technique. Table 2 displays the coefficients' significance of the GLMM for each selected covariance structure.

Table 2: *Estimates of the fixed effects of the GLMM model*

Parameter	Estimate	Std. Error	df	p-value	95% Confidence Interval	
Assuming AR(1) structure						
Intercept	1.216175	.932792	21.736	.206	-.719678	3.152028
TIME * COUNTRY	-.000268	.000256	2168.732	.296	-.000770	.000235
TIME	.004622	.003363	2168.732	.170	-.001974	.011217
COUNTRY	.106919	.071022	21.736	.147	-.040475	.254312
Assuming CS structure						
Intercept	1.417551	.963201	20.014	.157	-.591558	3.426660
TIME * COUNTRY	-9.04161E-005	2.55092E-005	2352.000	.000	-.000140	-4.03933E-005
TIME	.002073	.000335	2352.000	.000	.001416	.002730
COUNTRY	.102331	.073337	20.014	.178	-.050640	.255302
Assuming diagonal structure						
Intercept	1.417551	.963201	20.014	.157	-.591558	3.426660
TIME * COUNTRY	-9.04161E-005	2.55092E-005	2352.000	.000	-.000140	-4.03933E-005
TIME	.002073	.000335	2352.000	.000	.001416	.002730
COUNTRY	.102331	.073337	20.014	.178	-.050640	.255302

Dependent Variable: DAIRY PRICES.

Looking at Table 2, it is clear that when examining each item separately for the AR(1) variance-covariance within-country structure, the time effects on the dairy prices are not statistically significant. The same holds for the country effect included in the model as a covariate, as well as for their interaction. When considering that the real exchange rates (based on dairy prices) closer together in time should have higher correlations than observations that are further apart, then both effects of [TIME] and [COUNTRY] are no longer significant. Thus, the serial correlation between the measurements within the same country is the most dominant effect in the real exchange rates based on dairy prices data.

Regarding the GLMM with a CS variance-covariance structure, it is clear that not all effects are non-significant. Since the coefficient for time is positive (0.002), it may be concluded that time value is positively related to dairy prices. As regards the significant F-test for the time variable, it implies that dairy prices vary by year of measurement. As concerns the two-way interaction between the fixed factors, the *p*-value indicates the significance of the interaction (p-value<0.001), meaning that the [COUNTRY] effect differs as we move forward in time. On the other hand, the [COUNTRY] effect was found to be non-significant (p-value=0.178>0.05) at a 5% level of statistical significance.

The results of the fit of the last GLMM, assuming an independent variance-covariance structure, indicate that there are no significant effects of [year] (*p*-value=0.458) and its interaction with [COUNTRY] (*p*-value=0.657). On the other hand, as concerns the country effects, there appears to be a significant effect (beta coefficient=0.1018, *p*-value<0.001). The latter results indicate that ignoring the dependencies between the measurements within the same country could lead to entirely different results than taking into account the possible correlations.

The goodness-of-fit tests have shown that the best modelling approach for describing the associations between dairy prices and [COUNTRY], [TIME] effects is through the adoption of a mixed model, where observations on dairy prices within each country are assumed to be correlated. In particular, the autoregressive of order 1 [AR(1)] variance-covariance matrix structure is the structure that bests fits our data.

4.1. Unit Root – Cointegration tests results

Based on the results of the methodology mentioned above, the residuals that lack interdependencies were derived, and we employed the Johansen cointegration technique to examine whether there is a long-term relationship for the modified data. The unit root test preceded the application of the Johansen cointegration technique in order to confirm whether the time series employed are I (1). The results of the unit root tests (ADF test, GLS-DF test), presented in the following Table 3, offer no evidence as concerns the validity of PPP for the period studied. The only exception is the case of Lithuania for which both tests reject the null hypothesis of a unit root. Thus, for all the real exchange rates based on the dairy prices we conclude that they are not mean reverting for the period studied, which provided the background for employing the Johansen cointegration technique to survey whether the deviations from the PPP as expressed by real exchange rates are converging in the long-term.

Table 3: Results of Unit Root Tests (DF-GLS test, ADF test)

Country	DF-GLS test	ADF test	Country	DF-GLS test	ADF test
Austria	-0.653542	-1.100353	**Portugal**	-0.736030	-1.282009
Belgium	-0.894144	-1.698325	**Spain**	-0.8090466	-1.228588
Bulgaria	-2.671069	-1.983609	**Sweden**	-0.991164	-1.175333
Finland	-0.870462	-0.780508	**Denmark**	-1.197548	-0.866794
Germany	-0.904938	-0.984067	**Cyprus**	-1.160068	-1.197307
Estonia	-0.633959	-0.450058	**United Kingdom**	-1.360011	-0.912137
Latvia	-4.360880*	-4.653802*	**France**	-0.727949	-0.640668
Lithuania	-1.716910	-1.894705	**Italy**	-1.068686	-0.951828
Luxemburg	-0.5384	-1.276000	**Ireland**	-0.851266	-0.883503
Malta	-1.082205	-1.195867	**Greece**	-1.135710	-1.033826
The Netherlands	-0.946682	-1.665976	**Norway**	-1.085510	-1.134562

Notes: The optimal lag length is based on The Akaike criterion. * Indicates rejection of the null hypothesis at 5% level of significance. The critical values for 1,5 and 10% level of significance for the DF-GLS test and ADF test are respectively; -3.572800 ,-3.024000,-2.734000 and -4.046925, -3.452764, -3.151911.

Based on the results from the ADF test we estimated the half-life of the deviation from PPP, which is defined as the number of months it takes for deviations from PPP to subside permanently below 0.5 in response to a unit shock in the level of the series. The formula used for the calculation of the half-life is the following;

$$t_{1/2} = \frac{\ln(0.5)}{1+\delta}.$$

where t the number of months.

The implementation of the Johansen cointegration technique indicated the existence of long-term relationships for certain groups of countries implying homogeneity in the formation of consumer dairy prices or even different rates of pass through effects from producer to consumer prices. A significant and long-lasting food price pass-through effect has been identified and specifically, three groups were identified for which the existence of a sole cointegrating vector was confirmed.

The first group includes the following new member states of the EU: Bulgaria, Estonia, Latvia and Lithuania. The common trait for these countries is that they are economies in transition, while their economic growth is based significantly on agriculture. In most cases, prices were significantly below the EU level before accession. However, prices increased and converged after accession and almost reached the level set by the intervention price. The only exception is Hungary, which evidently is not included in our sample. The results confirmed the one sole relationship in the long-term that implies the existence of dependency among the dairy consumer prices for the aforementioned countries (Table 4).

Table 4: *Results of Johansen cointegration technique for Bulgaria, Estonia, Latvia and Lithuania*

Trace Test				
Hypothesized No. of CE(s)	**Eigenvalue**	**Trace Statistic**	**5 % Critical Value**	**1 % Critical Value**
None *	0.231119	53.00391	47.21	54.46
At most 1	0.141779	25.40795	29.68	35.65
At most 2	0.085149	9.354152	15.41	20.04
At most 3	9.26E-05	0.009728	3.76	6.65
Maximum Eigenvalue Test				
Hypothesized No. of CE(s)	**Eigenvalue**	**Max-Eigen**	**5 % Critical Value**	**1 % Critical Value**
None *	0.231119	27.59596	27.07	32.24
At most 1	0.141779	16.05380	20.97	25.52
At most 2	0.085149	9.344423	14.07	18.63
At most 3	9.26E-05	0.009728	3.76	6.65

The second group (Table 5) includes the South European countries of Cyprus, Greece, Italy, Malta, Portugal and Spain, which are characterized by a significant dairy product’s production.

Table 5 *Results of Johansen cointegration technique for Cyprus, Greece, Italy, Malta, Portugal and Spain*

Hypothesized No. of CE(s)	Eigenvalue	Trace Test Trace Statistic	5 % Critical Value	1 % Critical Value
None **	0.381557	124.5005	104.94	114.36
At most 1	0.253687	74.04276	77.74	85.78
At most 2	0.195484	43.31872	54.64	61.24
At most 3	0.112125	20.47968	34.55	40.49
At most 4	0.043396	7.992619	18.17	23.46
At most 5	0.031255	3.334178	3.74	6.40
		Maximum Eigenvalue test		
None **	0.381557	50.45777	42.48	48.17
At most 1	0.253687	30.72404	36.41	41.58
At most 2	0.195484	22.83904	30.33	35.68
At most 3	0.112125	12.48706	23.78	28.83
At most 4	0.043396	4.658441	16.87	21.47
At most 5	0.031255	3.334178	3.74	6.40

The last group involves traditional EU member states with different behaviour concerning the imports, the exports and the dairy product's production (Table 6). The results confirm the existence of a sole cointegrating vector and thus the existence of prices' interlinkages among them, since the real exchange rates converge in the long-term.

Table 6: *Results of Johansen cointegration technique for Finland, Netherland, Ireland, United Kingdom, and Germany*

Hypothesized No. of CE(s)	Eigenvalue	Trace test Trace Statistic	5 % Critical Value	1 % Critical Value
None *	0.368636	102.9097	94.15	103.18
At most 1	0.192815	54.62303	68.52	76.07
At most 2	0.120157	32.13178	47.21	54.46
At most 3	0.104184	18.69048	29.68	35.65
At most 4	0.065075	7.138350	15.41	20.04
At most 5	0.000695	0.073008	3.76	6.65
		Maximum Eigenvalue test		
None **	0.368636	48.28669	39.37	45.10
At most 1	0.192815	22.49125	33.46	38.77
At most 2	0.120157	13.44130	27.07	32.24
At most 3	0.104184	11.55213	20.97	25.52
At most 4	0.065075	7.065342	14.07	18.63
At most 5	0.000695	0.073008	3.76	6.65

The result is empirically confirmed, given that structural differences can be ignored due to the initial step of our methodology. Germany has high dairy prices as a large importer of dairy products, whereas the opposite holds for Ireland given that is one of the most significant intra - exporter. As for the Netherlands, they are a significant importer and exporter as well resulting in prices being at an intermediate level. Nevertheless, the main reason for the convergence of real exchange rates is the introduction of the CAP reform in 2003 through which the demand has started to be the driver for the market function.

Markets often deviate from this long-run equilibrium path due to various exogenous shocks and internal dynamism and reinstate the original long-run equilibrium path only when some error correction process begins; despite the existence of a long-run equilibrium confirmed with the implementation of cointegration technique, in the short-run. Consequently, the next step involved the estimation of the vector error correction model through which we estimated the speed of adjustment to the long-run equilibrium of the dairy consumer prices compared to the dairy prices of USA as expressed by the real exchange rates. The results for the three groups of countries are presented in the following Table 7.

Table 7: *Long-run elasticities and speed of adjustment for the first group of countries*

Error Correction:	**Bulgaria**	**Estonia**	**Latvia**	**Lithuania**		
Cointegrating Vector	1.0000	-5.829564	27.86864	4.124342		
		[-1.68778]	[4.79820]	[1.35263]		
Speed of adjustment	-0.060104	0.006042	-0.005209	-0.003467		
	[-1.88177]	[2.56321]	[-3.41007]	[-1.24848]		
Error Correction:	**Cyprus**	**Greece**	**Italy**	**Malta**	**Portugal**	**Spain**
Cointegrating Vector	1.000000	3.829635	-2.321227	2.098109	1.117472	-1.762176
		[4.90830]	[-4.03706]	[4.49587]	[1.49863]	[-3.80999]
Speed of adjustment	-0.018493	-0.303744	-0.448775	-0.089123	-0.020663	-0.021915
	[-0.28976]	[-5.56916]	[-4.47552]	[-1.94256]	[-0.39102]	[-0.39275]
Error Correction:	**Austria**	**United Kingdom**	**Germany**	**Ireland**	**Luxemburg**	**The Netherlands**
Cointegrating Vector	1.000000	-1.263001	-0.515713	1.114776	4.221611	-2.161496
		[-1.38687]	[-0.41941]	[2.44102]	[2.61863]	[-1.64199]
Speed of adjustment	-0.021489	-0.000872	0.015996	-0.003235	-0.026844	-0.033296
	[-1.89984]	[-0.07237]	[0.86630]	[-0.09515]	[-1.91741]	[-1.78127]

The long-run dynamics that denote the speed of adjustment to long-run equilibrium, suggest that the real exchange rates based on dairy consumer prices respond to deviations from long-run equilibrium, while the negative sign of the coefficient, implies that every shock in the real exchange rate based on the dairy prices of the group is absorbed with time and thus it tends to be eliminated, with a prolonged rate though. The dairy markets of Bulgaria, Estonia and Latvia are statistically insignificantly different from zero at a 5% significance level, while the opposite is confirmed for Lithuania.

This indicates that the real exchange rates for these are weakly exogenous and it is possible that they will not change in response to deviations from the long-run equilibrium. Additionally, the real exchange rates for Latvia and Estonia are profoundly affected by changes in those for Bulgaria and Lithuania, while the opposite is valid for Bulgaria and Lithuania in price changes of Latvia and Estonia. Finally, the speed of adjustment is almost equal in size, yet opposite in sign. The result denotes that the system would respond with a decrease for Latvia and an increase for Estonia in a positive deviation for the real exchange rates.

Similar conclusions can be reached for the second group that involves countries with particularities in their market conditions. The dairy market of Cyprus, Portugal and Spain are statistically non-significantly different from zero at 5% significance level, while the opposite is confirmed for Greece, Italy and Malta. This result indicates that the real exchange rates in Cyprus, Portugal and Spain are weakly exogenous and thus it is more likely that they will not change in response to deviations from the long-run equilibrium. Furthermore, a shock in the dairy markets (dairy prices) of Cyprus, Portugal and Spain affect the dairy markets of Greece, Italy and Malta strongly. On the other hand, a change in the real exchange rates of Greece, Italy and Malta has a small impact on the real exchange rates in Cyprus, Portugal and Spain. Finally, a comparison in the speed of adjustment implies that Malta needs the longest adjustment time to the equilibrium level relative to Italy and Greece.

Regarding the third group of countries, the real exchange rates for the United Kingdom, Germany and Ireland are weakly exogenous given that the speed of adjustment is statistically insignificant from zero for 5% level of significance, and thus the real exchange rates and, consequently, the milk prices in these markets may not change in response to deviations from the long-run equilibrium. On the contrary, Austria, Luxemburg and the Netherlands exhibit a speed of adjustment that is statistically significant from zero for 5% level of significance. This indicates that the real exchange rates for the United Kingdom, Germany and Ireland are less affected by changes in the prices of Austria, Luxemburg and Netherlands. Regarding the speed of adjustment, the system would respond with a decrease in the real exchange rates for Austria, Luxemburg and the Netherlands gave a positive deviation from the long-run equilibrium. It is worth mentioning that comparing the speed of adjustment for those three countries, they adjust almost simultaneously at the steady state given a change, for example, in Germany or the United Kingdom

5. Conclusions

This paper set out to investigate the existence of an integrated European dairy market through the application of mixed model effects to survey the existence of serial correlation within the time series studied (among the observations). The estimation of the mixed models generated residuals free from serial correlation on which the Johansen cointegration technique was subsequently employed. The implementation of the particular technique identified three significant groups of dairy markets for which a single cointegration vector was confirmed.

The results validated the existence of a long-run relationship among real exchange rates based on dairy prices for these groups. Each group of EU countries is characterized by a specific feature, which in the case of the first group, is that all group members are transition economies that have recently entered the EU. The second group includes South European countries with, actually, complicated producer-consumer chains, while the third group involves old EU member states with well-organised systems to provide the consumers with dairy products. Weakly exogenous variables existed across all groups, offering the other countries, within each group, protection against external shocks and adversaries of price volatility.

As concerns the markets of each identified group, there seems to be integrated into the long run, given the single cointegrating vector that links them, while the presence of the weakly exogenous variables indicates that under the current circumstances, they are sending low price signals to other markets. This may imply that the markets, in which the real exchange rates based on dairy prices are exogenous variables, will not be able to influence each other’s prices and to some extent, they behave independently in the short-run. If this will be the case, it will restrain the progress and pace of effectiveness to many national milk market policies and, in the short-run, a macro policy implemented by EU may not yield the desirable and uniform results within the EU.

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