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Price linkages in the international skim milk powder market: empirical evidence from nonparametric and time-varying copulas

Panos Fousekis, Christos Emmanouilides and Vasilis Grigoriadis[†]

The objective of this article is to assess the integration of the international skim milk powder (SMP) market. This is pursued using monthly data over 2001 to 2014 from the three principal SMP-producing regions (the EU, the USA and Oceania) and nonparametric kernel-based and time-varying copulas. The empirical results point to a strong and an increasing degree of overall price co-movement and to statistically significant probabilities for joint price crashes and booms. While the EU and Oceania have been the regions with the highest degree of integration, the USA has been catching up with them.

Key words: kernel-based copulas, price co-movement, skim milk powder.

1. Introduction

Spatial price relationships has been an important topic in theoretical and empirical research for a long period of time. The keen interest of economists lies in the recognition that the strength and the pattern of price linkages provide information on whether spatial markets are integrated (globalised) or segmented (regionalised). In well-integrated markets, prices tend to co-move. Transmission of price shocks from one spatial market to another is a necessary condition for economic efficiency (Ghoshray 2010; Reboredo 2011; Emmanouilides and Fousekis 2012). The degree of integration, in turn, has implications about the level of the total welfare obtained as well as about the distribution of benefits among the markets (regions) involved in trade (Meyer and von Cramon-Taubadel 2004; Serra *et al.* 2006).

The empirical assessment of spatial price linkages has been carried out with a large variety of econometric approaches, both parametric and nonparametric. The former includes linear integration and cointegration (Dawson and Dey 2002), and nonlinear integration and cointegration (Goodwin and Piggott 2001; Abdulai 2000; Emmanouilides and Fousekis 2012). The latter includes local polynomial regression (Serra *et al.* 2006). Emphasis has been

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placed not only on the strength of price linkages but also on the identification and quantification of other potentially relevant characteristics such as asymmetric price transmission (Goodwin and Piggott 2001; Serra *et al.* 2006; Reboredo 2011).

Against this background, the objective of this article is to investigate price linkages in the international skim milk powder (SMP) market. The international trade of dairy products has gained a considerable momentum in recent years due to: (i) food trade liberalisation through multilateral and bilateral agreements, (ii) innovations in milk processing and (iii) rising per capita incomes, changing demographics and Westernisation of diets in a number of countries facing natural handicaps in dairy production because of climate and/or land and feed scarcity (Beghin 2005; Dong 2006). The SMP is the third most important, in terms of product weight, internationally traded dairy commodity (after whole milk powder and cheese). The SMP market, however, is the most competitive among all international dairy commodity markets (AgriHQ Academy, 2014).

The assessment of spatial price linkages in the international SMP market is pursued here using *copulas*, a statistical tool which recently has become increasingly popular in modelling co-movement (dependence) among random variables. Copulas have been used extensively in engineering, finance and risk management. There have been, however, very few published empirical works with copulas on market integration either in the physical or in the product quality space. Reboredo (2011) examined price co-movement in regional crude oil markets. Serra and Gil (2012) analysed the linkages between biodiesel, diesel and crude oil prices in Spain. Qiu and Goodwin (2013) and Emmanouilides and Fousekis (2015) evaluated price transmission along the pork and the beef, respectively, supply chains in the US. Emmanouilides *et al.* (2014) assessed price dependence in major olive oil markets of the Mediterranean, and Zimmer (2015) investigated price co-movement between corn, soya bean and wheat in the USA.

The above works on market integration with copulas have adopted the parametric approach in which a researcher either specifies a parametric copula family *a priori* on the basis of its 'desirable' properties or, at best, she (he) selects formally among a few parametric families using certain information criteria. Given that (i) the number of parametric copula families available in the literature is very large and (ii) a copula family which outperforms a few other families is not necessarily the most appropriate representation of the joint stochastic process under study, the parametric approach may result in misspecification. This potential problem may be further exacerbated when co-movement is time-varying since the solutions offered in that case such as modelling one or more dependence parameters as autoregressive processes are largely *ad hoc*.¹

¹ For details on modelling time-varying parametric copulas, see Patton (2013) and Qiu and Goodwin (2013).

The present work contributes to the literature of spatial price transmission and market integration for agricultural and food commodities from at least two viewpoints. First, it focusses on an important international market which has exhibited strong growth in recent years and is expected to expand further in the future. We are not aware of any earlier empirical works on price inter-relationships in the international markets of dairy commodities, in general. Second, for the investigation of price co-movement, it relies on the kernel-based nonparametric approach (Fermanian and Scaillet 2003; Racine 2015) which provides a smooth (differentiable) reconstruction of the copula function dispensing with the need to specify a mathematical form. To the best of our knowledge, there have been no earlier published empirical studies on spatial market integration with kernel-based copulas. Further, it employs a nonparametric procedure to test for time-varying co-movement and a moving window to trace out how price linkages in the international SMP market have evolved over time.

The empirical analysis here utilises prices from the three top producers and exporters of SMP (the EU, the Oceania and the USA). Although Oceania is a small economic region relative to the EU and to the USA, it is a major player in the international SMP markets contributing one-fifth of the global production and one-third of the global exports. According to the results, the degree of market integration for the three regions has been considerable over 2001 to 2014, and more importantly, it has been steadily rising over time. The increase in global demand for SMP, the decision by the European Commission to reduce the intervention price of the commodity and the emergence of the USA as a major player in the international markets are likely to be behind these developments. In what follows, Section 2 presents the analytical framework and Section 3 the data, the empirical models and the results. Section 4 offers conclusions and suggestions for future research.

2. Analytical framework

The use of copulas to assess co-movement among random processes has its root in Sklar's (1959) theorem according to which a multivariate distribution function is fully specified by its univariate marginal distributions and a joining function termed as *copula*. For the bivariate case, let $Y = (Y_1, Y_2)'$ be a random vector with joint distribution function F and with marginal distribution functions $F_i(y_i)$, $i = 1, 2$. Then, F can be expressed as

$$F(y) = C(F_1(y_1), F_2(y_2)) \quad (1)$$

where C is the copula function mapping the univariate marginal distributions to the joint distribution. Provided that a marginal distribution is continuous, its *probability integral transform* $U_i = F_i(y_i)$ follows the uniform distribution

on $[0,1]$. Consequently, the copula function may be viewed as the joint distribution function of the probability integral transforms, that is, one with uniform marginals on $[0,1]$ ($C : [0,1]^2 \rightarrow [0,1]$).

In $C(F_1(y_1), F_2(y_2))$, each element of Y is fed into its own distribution function. As a result, all information contained in the marginal distributions is swept away and what is left in C is the pure joint information between Y_1 and Y_2 (Meucci 2011). By capturing the information missing from the marginal distribution to complete the joint distribution, the copula characterises fully and uniquely co-movement of the elements of the random vector (Trivedi and Zimmer 2005; Emmanouilides and Fousekis 2015).

A substantial advantage of copulas is that they allow one to specify marginal distributions separately from their dependence structure (copula); Nelsen 2006; Patton, 2013. In other words, in copula modelling, marginal distributions and dependence are treated as two separate but related issues (Trivedi and Zimmer 2005) offering, thus, considerable flexibility in empirical research. Patton (2013) notes that a researcher may combine a normally distributed random variable with an exponentially distributed random variable via a Student- t copula and obtain a valid bivariate distribution. A researcher, therefore, may draw on the very large number of univariate distributions leaving only the task of modelling their dependence structure.

A key attribute of any co-movement (dependence) measure is *scale invariance* (meaning invariance to strictly increasing transformations of the underlying stochastic process). Scale invariant dependence measures are functions of the ranks (or the probability integral transforms) of the data only. Equivalently, scale invariant (or *pure*) dependence measures are functions solely of the copula and not of the marginal distributions (Nelsen 2006; Patton 2013).

A number of pure measures of *global (overall)* bivariate co-movement (e.g. Kendall's *tau*, Spearman's *rho*, Bloomqvist's *beta*, Hoeffding's *phi*, and the Gini's *gamma*) have been long used in economic and social research. In this work, we consider Hoeffding's *phi* defined as (Schmid *et al.* 2010)

$$\Phi(C) = \left(\frac{1}{90} \int_{[0,1]^2} (C(U) - \Pi(U))^2 dU \right)^{\frac{1}{2}}. \quad (2)$$

Φ may be interpreted as the (normalised) distance between a given copula C and the independence copula Π . Its range is $[0,1]$, receiving the value of zero when $C = \Pi$ and the value of one when $C = M$ or $C = W$, where M and W are the upper and lower Fréchet-Hoeffding bounds of copula C , respectively. For two-dimensional random vectors, both M and W are copulas, with M corresponding to co-monotonicity and W to counter-monotonicity of the

random variables (Schmid *et al.* 2010).² Hoeffding's *phi* is our preferred measure of global co-movement because it has better properties relative to other commonly used measures (Gaißer *et al.* 2010; Schmid *et al.* 2010).

The measures of global (overall) co-movement summarise the strength/intensity of dependence between random processes over their entire support. However, the intensity of dependence is not necessarily the same at all parts of a joint distribution (e.g. at the centre and at the tails). Richer insights about the relationships under study and probably more relevant information (the economic agents are more likely to be interested in the strength of the price dependence during extreme market events rather than under during less drastic changes) may be obtained from the analysis of *local co-movement*, that means, the co-movement over parts of the support. The relevant notions for assessing local co-movement are the *quantile dependence coefficients*. These are given by

$$\lambda_L^q = P(U_2 \leq q / U_1 \leq q) = \frac{C(q, q)}{q}, \quad 0 < q \leq \frac{1}{2}, \quad (3)$$

$$\lambda_U^q = P(U_2 > q / U_1 > q) = \frac{1 - 2q + C(q, q)}{1 - q}, \quad \frac{1}{2} < q < 1. \quad (4)$$

Equation (3) defines a set of *lower quantile dependence coefficients* providing the conditional probability that the random process Y_2 receives a value at most equal to its q quantile given that the random process Y_1 receives a value at most equal to its q quantile, as well. Equation (4) defines a set of *upper quantile dependence coefficients* providing the conditional probability that the random process Y_2 receives a value strictly higher than its q quantile given that the random process Y_1 receives a value strictly higher than its q quantile, as well (Nelsen 2006; Patton 2013). By varying q , one may trace out how the intensity of relationship behaves at the different parts of the support.

For the analysis of spatial market integration of particular importance is the co-movement of prices at the extreme lower and at the extreme upper quantiles. As noted by Reboredo (2011) and Qiu and Goodwin (2013), joint booms and joint crashes of prices indicate well-integrated spatial markets.³ Technically, this implies sizable and statistically significant dependence coefficients at the very high and the very low quantiles of the joint

² Under stochastic independence, $C(U_1, U_2) = U_1 U_2 = \Pi(U_1, U_2)$ and Hoeffding's *phi* equals zero. The upper and the lower Fréchet-Hoeffding bounds specify the range of values a copula may achieve. In particular, for every pair (U_1, U_2) in the domain of a copula C it is the case that $W = \max(U_1 + U_2 - 1, 0) \leq C(U_1, U_2) \leq \min(U_1, U_2) = M$ (Nelsen 2006; Repkine, 2014). The term co-monotonicity refers to perfect positive co-movement while the term counter-monotonicity refers to perfect negative co-movement.

³ Another important indicator is, of course, the degree of overall co-movement.

distribution. Using the quantile dependence coefficients, one may also formally test for asymmetric co-movement (e.g. whether the intensity for $q \leq 0.2$ is the same as that for $q > 0.8$). The quantile dependence coefficients, therefore, provide information both about the strength as well as about the structure of co-movement.

3. The data, the empirical models and the results

The data for the empirical investigation are monthly wholesale prices of SMP (in \$1000/tonne). They have been obtained from DairyCo and they refer to the period 2001:1 to 2014:12.⁴ The EU, the USA and Oceania are, in this order, the three major SMP producers at a global level. They are also the three major exporters. Taken together, they accounted for more than 73 per cent of the global production and for more than 82 per cent of the global exports in 2011–13 (Appendix S1). The principal import markets for SMP are Mexico, China, South East Asia, Russian Federation, Middle East and North Africa.

The SMP is considered to be the ideal dairy commodity for emerging markets with lower purchasing power. It is used for the recombination of milk products and for other applications in the food industry as well. Although demand for SMP in developed countries has been stagnating in recent years, growth in SMP demand has risen from developing countries where not enough freshly produced milk is available. Trade flows of SMP among the USA, the EU and Oceania are negligible (if any). Integration of the three spatial markets under these circumstances may be achieved indirectly (i.e. through their competition on the international SMP market).

Appendix S4 presents the natural logarithms of SMP prices. The series tend generally to move together providing, thus, an indication that price shocks are transmitted from one SMP producing region to the other(s). On average, the USA has been the region with the lowest price (\$2627/tonne) and Oceania has been the region with the highest price (\$2827/tonne).

Following earlier empirical works on market integration with copulas (Reboredo 2011; Serra and Gil 2012; Emmanouilides and Fousekis 2015), this article focuses its attention on the co-movement among the rates of price change (price shocks), calculated as $d \ln p_{it}$, where p_{it} is the price of SMP in region i = the USA, Oceania and the EU at time $t = 1, 2, \dots, T$; Appendix S2 presents summary statistics for the price shocks in the three regions.

The asymptotic properties of different copula estimators have been established under the assumption of *i.i.d.* observations (Fermanian and Scaillet 2003; Patton 2013). Time series data, however, may exhibit autocorrelation and ARCH effects. To address this potential problem, following Reboredo (2011) and Serra and Gil (2012), we have filtered the

⁴ www.dairyco.org.uk/market-information. The choice of the time period has been constrained by the data availability.

individual series of raw price shocks using skewed- t ARMA-GARCH marginal models (Appendix S3). Filho *et al.* (2012) note that skewed versions of GARCH models ensure that any asymmetry found for a multivariate covement structure is genuine and not a consequence of marginal misspecification. The asymmetry coefficients turned out to be positive and statistically significant for all estimated models indicating long upper tails of the respective marginal distributions (i.e. large positive shocks, as observed in upturn periods, have been more likely than large negative shocks, as observed in downturn periods). The shape coefficients turned out to be positive and statistically significant indicating leptokurtic (heavy-tailed) marginal distributions. The estimated degrees of freedom have been very low (ranging from 2.4 to 3.4) also suggesting strong departure of the raw price shocks from normality. Table 1 presents the P -values from the application of the Box-Pierce and the ARCH-LM tests to the standardised innovations (*filtered price shocks*) at a number of lag lengths. It appears that the filtered data are free from autocorrelation and ARCH effects.

A copula is a distribution function with uniform margins on $[0,1]$. For the empirical application, therefore, the standardised innovations have been converted to *copula data* using probability integral transforms and a scaling factor equal to $T/T + 1$, where T is the number of observations (Serra and Gil 2012; Emmanouilides and Fousekis 2015). The nonparametric Kolmogorov–Smirnov (KS) goodness of fit test has been, then, applied to each individual copula data series to verify whether it actually follows the uniform distribution on $[0,1]$. Table 2 presents the results. In all cases, the null hypothesis that the empirical distribution is consistent with the hypothesised distribution (uniform on $[0,1]$) has not been rejected at any reasonable level of significance.

To assess whether each of the three bivariate copulas, (USA and Oceania), (USA and EU) and (Oceania and EU), is time invariant, we have employed the nonparametric test by Buseti and Harvey (2011) which is based on sample q -*biquantics* (Appendix S5). Following Buseti and Harvey (2011) and Emmanouilides and Fousekis (2015), we have calculated empirical values of the relevant test statistic at three quantiles, namely 0.25, 0.50 and 0.75. Table 3 presents the results. For the copula (USA, Oceania), the null has

Table 1 P -values of the tests for autocorrelation and ARCH effects (filtered price shocks)

Standardised innovations	Box-Pierce			ARCH-LM		
	No. of lags			No. of lags		
	1	6	12	1	6	12
USA	0.117	0.727	0.810	0.977	0.929	0.911
Oceania	0.232	0.685	0.105	0.115	0.548	0.912
EU	0.195	0.624	0.619	0.353	0.731	0.771

Table 2 Results from the application of the Kolmogorov–Smirnov (KS) test to the copula data

	Empirical value of the KS statistic†
USA	0.072 (0.782)
Oceania	0.054 (0.969)
EU	0.048 (0.991)

Note P -values in parentheses. †The test statistic is obtained as the $\sup_y |F_{\text{theo}}(y) - F_{\text{emp}}(y)|$, where y is the data set, F_{theo} is the theoretical distribution (here, Uniform on $[0,1]$) and F_{emp} is the empirical distribution.

Table 3 Results from the application of the Buseti–Harvey test to the bivariate empirical copulas†

Empirical copula	Quantiles		
	0.25	0.50	0.75
(USA, Oceania)	0.661	0.387	0.170
(USA, EU)	0.096	0.517	0.381
(Oceania, EU)	0.416	0.374	0.184

Note †The critical values are 0.743, 0.461 and 0.347 at the 1, 5 and 10 per cent level, respectively.

been rejected at the 10 per cent level for the 0.5 quantile and at the 5 per cent level for the 0.25 quantile. For the copula (USA, EU), the null has been rejected at the 10 per cent level for the 0.75 quantile and at the 5 per cent level for the 0.5 quantile. For the copula (Oceania, EU), the null has been rejected at the 10 per cent level for the 0.25 and the 0.5 quantiles. Overall, the empirical evidence suggests that price dependence has not been time-invariant over the period 2001–2014.

To trace out how price co-movement has evolved over time, we have estimated the three bivariate copula functions for a number of subperiods in 2001–2014 using a moving window of length 96 months (8 years). The choice of the length of the subperiod constitutes a compromise between the need to base each subperiod estimation on an adequate number of observations and the need to obtain sufficient detail on the dynamics of co-movement. With that window length and with 167 observations available, the number of subperiods is 72. Appendix S6 provides information on the computation of Hoeffding's ϕ from the copula data.

Figure 1 (panels a–c) presents the estimates of global co-movement along with their 95 per cent confidence bands for each of the 72 subperiods.⁵ For all three pairs of rates of price change and for all subperiods considered, Hoeffding's ϕ is statistically significant at the 5 per cent level or less. For the pair USA and Oceania, Hoeffding's ϕ almost doubled from about 0.25 in the early subperiods to about 0.5 in recent ones. For the pair USA and EU, it has risen from about 0.35 in the early subperiods to about 0.55 in recent ones.

⁵ The confidence bands come from the application of block bootstrap (Politis and Romano 1994; Patton *et al.* 2009; Patton 2013), with 500 replications.

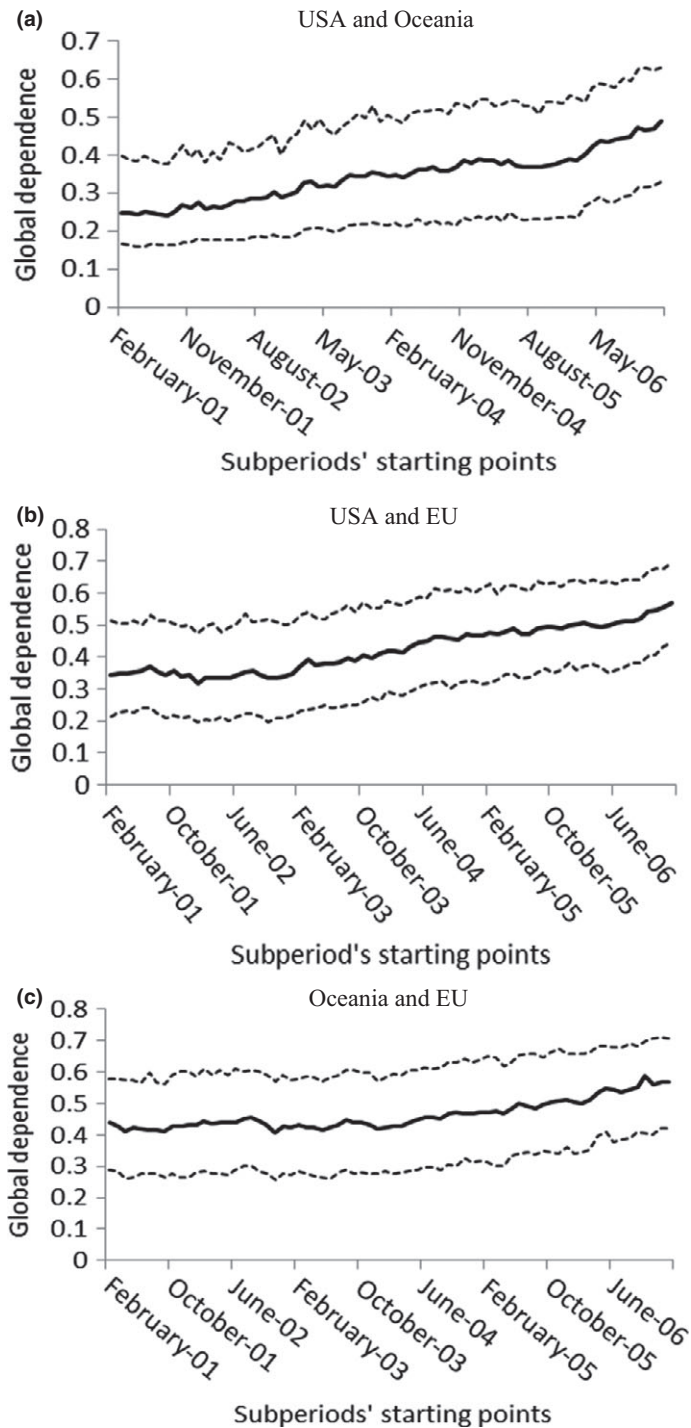


Figure 1 Hoeffding's Φ and 95 per cent confidence bands.

For the pair Oceania and EU, Hoeffding's *phi* has risen from about 0.42 in the early subperiods to about 0.56 in recent ones. The analysis of global co-movement, therefore, suggests that (i) Oceania and the EU have been the regions with the highest degree of integration, (ii) the USA started with a considerably lower degree of integration relative to the other two regions but it has been closing the gap at a very fast pace and (iii) the interconnection between all three major SMP-producing regions has been increasing over time. These results appear to be in line with the developments in the international SMP market in the last 15 years. Global trade of SMP has doubled from <1 million tonnes in 2000–2002 to 1.8 million tonnes in 2011–2013. The intensification of trade has facilitated spatial price transmission. The USA (thanks to favourable international prices) has become very active in the international SMP market, especially after 2007/2008. It has increased export volumes and it has, to a certain extent, differentiated its export markets. While Mexico still remains the major importer of the US-produced SMP, markets in Asia, North Africa and Middle East have been gaining importance. On those latter markets, the USA exporters have to compete for customers with their counterparts from Oceania and from the EU. The prices in Oceania and in the EU exhibit higher degree of overall co-movement because these two regions have been the main competitors in the international SMP markets over the sample period. Price transmission between Oceania and the EU has been further enhanced by the gradual reduction in the EU intervention price for SMP since the early 2000s (Thiele *et al.* 2013).

The kernel-based nonparametric estimation of the copula functions has been implemented here using the '*inversion approach*' (Fermanian and Scaillet 2003; Balakrishnan and Lai 2009; Racine 2015). In particular, solving Equation (1) obtains

$$C(U_1, U_2) = F(F_1^{-1}(U_1), F_2^{-1}(U_2)) \quad (5)$$

where $F_i^{-1}(U_i) = \inf(y/F_i(x) \geq U_i)$ ($i = 1, 2$) is the *quasi-inverse* of F_i . A kernel-based estimator of F_i is

$$\hat{F}_i = T^{-1} \sum_{t=1}^T K_i \left(\frac{y_i - Y_{it}}{h_i} \right) \quad (6)$$

and a kernel-based estimator of F is

$$\hat{F} = T^{-1} \sum_{t=1}^T \prod_{i=1}^2 K_i \left(\frac{y_i - Y_{it}}{h_i} \right), \quad (7)$$

where h_i is the bandwidth and K_i is the cumulative distribution of an appropriate two-dimensional kernel function. The copula estimate at point u is then $\hat{C}(U) = \hat{F}(\hat{\zeta}_1, \hat{\zeta}_2)$, where $\hat{\zeta}_i = \inf(y/\hat{F}_i(y) \geq U_i)$ is the kernel estimate of

the Y_{it} 's quantile with probability level U_i . The basic idea of kernel-based nonparametric estimation is to 'let the data speak for themselves' rather than to impose *a priori* restrictions (through a functional form) on the strength and the pattern of co-movement. Therefore, the nonparametric estimation is generally more flexible than the parametric one.

Quantile dependence is measured by replacing the copula functions in (3) and (4) with their respective estimates.⁶ Among the potentially infinite number of such coefficients, considerable attention in the relevant literature has focused on the so-called *tail dependence coefficients* measuring co-movement at the very extremes of a joint distribution. Technically, the lower tail dependence coefficient is the limit a lower dependence coefficient approaches as q goes to zero, while the limit of an upper tail dependence coefficient is the value approached as q goes to one.

For bivariate parametric copula families, explicit mathematical expressions for the tail dependence coefficients are available. Their calculation, however, from nonparametric copulas presents a difficulty since if one sets q close enough to 0 (1) it causes the sample estimate of λ_L^q (λ_U^q) to go to zero as well (Patton 2013). As shown by Dobric and Schmid (2005), a 'cut-off' quantile in the neighbour of $1/\sqrt{T}$, where T is the sample size, provides a reasonable solution to the problem in the sense that the nonparametric estimates at the $1/\sqrt{T}$ and the $1 - 1/\sqrt{T}$ quantiles may be viewed as approximations to the lower and the upper tail dependence coefficients, respectively.

Here, with 96 observations available for each subperiod, we have selected the quantiles 0.10 and 0.90 to evaluate co-movement under extreme negative and extreme positive prices shocks, respectively, and the quantiles 0.40 and 0.60 to evaluate co-movement under general (both strong and weak) negative and positive price shocks, respectively. Figures 2–4 (panels a and b) present the estimates of the 0.10 and the 0.90 quantile dependence coefficients along with their 95 per cent confidence bands from the three bivariate copulas. The 0.10 and the 0.90 dependence coefficients are statistically significant at the 5 per cent level or less for all subperiods and for all kernel-based copulas suggesting that extreme positive and extreme negative price shocks are transmitted from one SMP-producing region to another. Co-movement at the extremes has been generally rising over time it has been consistently stronger for Oceania and the EU and weaker for the USA and Oceania. For example, in the most recent subperiods, an extreme positive price shock has been transmitted with a probability of about 0.5 for Oceania and the EU and with a probability of about only 0.35 for the USA and Oceania. The results from the analysis of co-movement at the extremes are, therefore, in line with what has been already transpired from the preceding analysis of Hoeffding's ϕ .

⁶ We have carried out the nonparametric kernel-based estimation of the copula functions using the np package in R (Version 0.60-2; R Development Core Team, 2014), multivariate least-squares cross-validation and Gaussian kernel functions (Li and Racine 2013; Racine 2015).

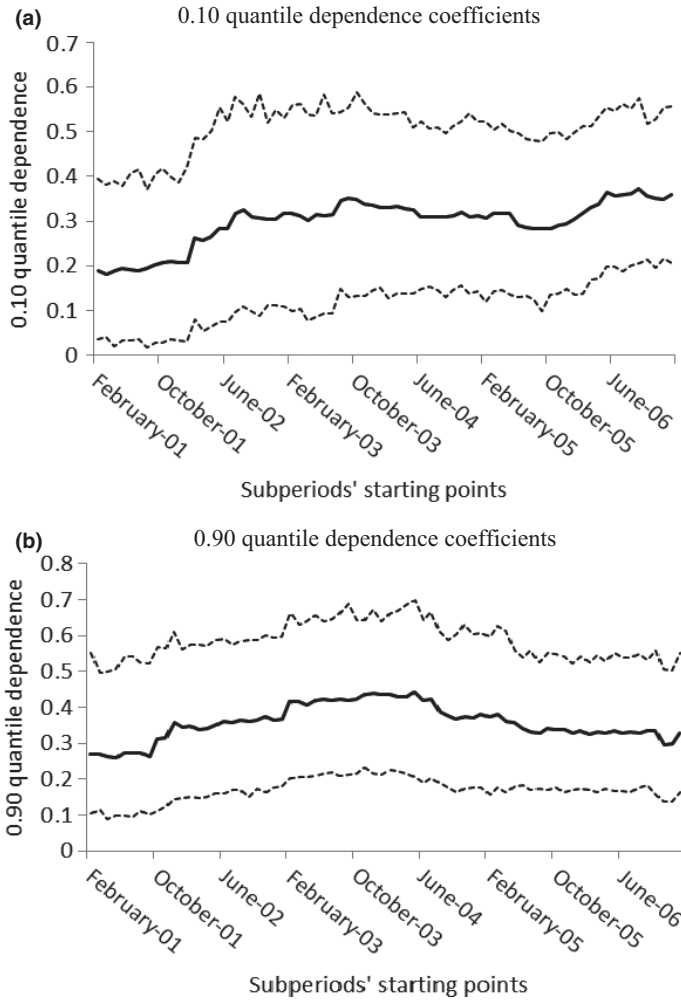


Figure 2 USA and Oceania. Quantile dependence coefficients at the 0.10 and the 0.90 and their respective 95 per cent confidence bands.

To avoid a clutter with figures, we have chosen not to present details on co-movement at the 0.4 and the 0.6 quantiles. The relevant point estimates and their respective 95 per cent confidence bands are available upon request. The findings, however, are qualitatively very similar. There have been upward trends in quantile dependence, especially for the USA and Oceania, and for the USA and the EU. In the most recent subperiods, the dependence coefficients at the 0.4 and at the 0.6 quantiles for all bivariate copulas lie in the range of 0.55–0.65.

The comparison of two upper (or two lower) quantile dependence coefficients provides information on the intensity by which price shocks of the same sign but of different magnitude are transmitted from one region to the other. To be specific, let us compare the values of the 0.10 and of the 0.40 quantile dependence coefficients for Oceania and the EU and for the last

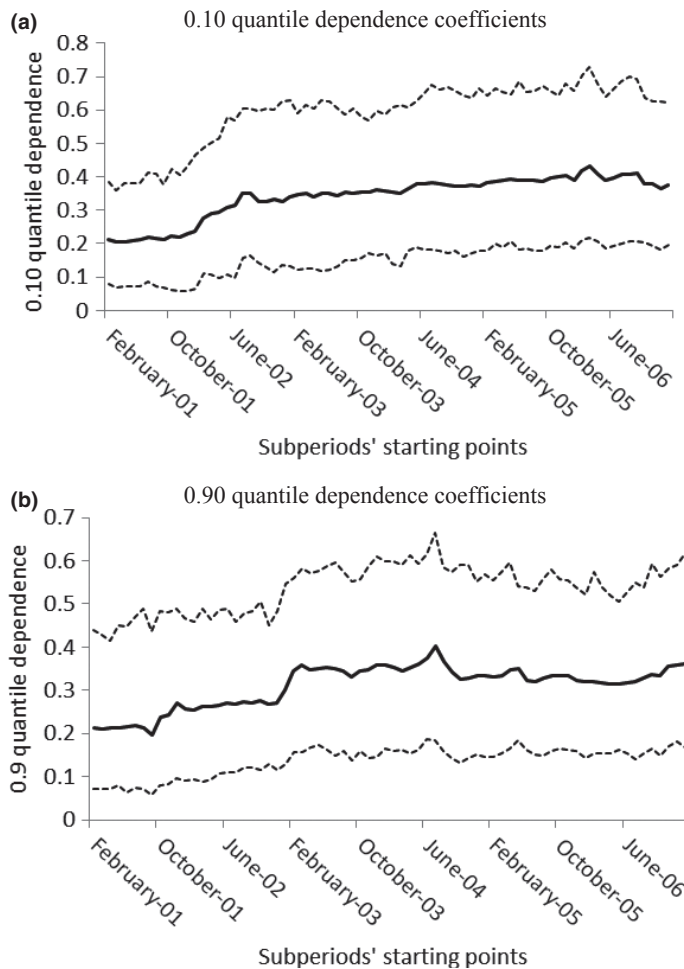


Figure 3 USA and EU. Quantile dependence coefficients at the 0.10 and the 0.90 and their respective 95 per cent confidence bands.

subperiod in the sample. These are 0.296 and 0.595, respectively, suggesting that the probability mass in the line interval $[0,0.1]$ is almost the same as the probability mass in the line interval $[0.1,0.4]$ (that is $0.299 = 0.595 - 0.296$). Given that the latter interval is three times as wide as the former, the result implies that the probability mass of the respective conditional distribution gets thinner as one moves closer to the median (responses diminish as price shocks become progressively smaller in absolute terms). In other words, it implies that smaller price shocks are less likely to be transmitted from one SMP-producing region to another compared to larger price shocks. This finding (which holds for all quantile dependence coefficients, all bivariate copulas and all subperiods considered in this work) is certainly consistent with the literature of spatial price transmission which points to inactivity bands in the neighbourhood of medians and to thresholds that price shocks in one region have to surpass in order to trigger responses to other regions. The

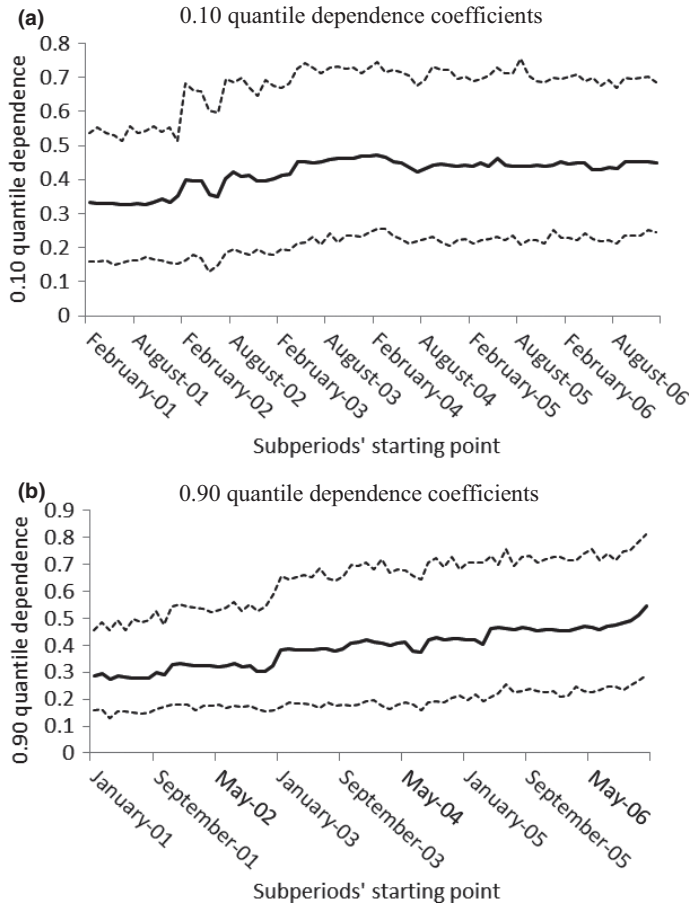


Figure 4 Oceania and EU. Quantile dependence coefficients at the 0.10 and the 0.90 and their respective 95 per cent confidence bands.

inertia/inactivity in those parts of the joint support is typically attributed to the presence of transaction costs (Obstfeld and Taylor 1997; Meyer and von Cramon-Taubadel 2004; Serra *et al.* 2006). It appears, therefore, that the statistical tool of copulas corroborates the empirical evidence obtained in earlier works from econometric models.

An important issue in market integration studies is that of asymmetric price transmission between shocks of opposite sign but of the same absolute magnitude (Goodwin and Piggott 2001; Reboredo 2011; Emmanoulides *et al.*, 2014). Here, to examine the presence of such asymmetries, we have conducted a Wald test (Patton 2013), the null hypothesis of which (symmetry) involves $\lambda_L^q = \lambda_U^{1-q}$, for $q < 0.5$. The relevant test statistic is

$$(\hat{\lambda}_L^q - \hat{\lambda}_U^{1-q})'(R\hat{V}_B R')^{-1}(\hat{\lambda}_L^q - \hat{\lambda}_U^{1-q}) \sim \chi_1^2, \quad (8)$$

where R is the restrictions matrix and \hat{V}_B is the bootstrap estimate of the variance–covariance matrix of the quantile dependence coefficients. For the

quantiles 0.10 and 0.90, the P -values associated with the above test statistic for all three bivariate copulas pairs and for all subperiods range between 0.32 and 0.99. Therefore, the null of symmetry cannot be rejected (extreme positive and extreme negative shocks appear to be transmitted with the same intensity). For the quantiles 0.40 and 0.60 in only 4 out 72 subperiods, the null has been rejected at the 10 per cent level or less. The lowest P -value for all three bivariate copulas and for all subperiods is 0.082. Again, the empirical evidence is overwhelmingly in favour of symmetric transmission of price shocks (the reaction of economic agents to 'bad news'/price downswings does not differ statistically from their reaction to the 'good news'/price upswings). A possible explanation behind the finding of symmetry is that milk powder, in general, and SMP, in particular, are sold on international markets as bulk (undifferentiated) commodities. Exporters, therefore, have a strong incentive to meet price changes initiated by their competitors.

The quantile dependence coefficients provide the probability of transmission of price shocks from one region to another but are not informative about the shocks' origins. To shed some light on this issue, we have performed Granger causality tests using the filtered data (pure innovations). Table 4 presents the results. There is strong evidence that the casual order flows unidirectional from Oceania to the USA and from the EU to the USA. For the pair EU and Oceania, the null hypothesis that price shocks from the former do not Granger-cause price shocks in the latter has been rejected at the 7.9 per cent level while the hypothesis that price shocks from Oceania do not Granger-cause price shocks in the EU has been rejected at the 9.6 per cent level. It appears, therefore, that EU and Oceania (the most export-oriented SMP-producing regions) are the causal regions. Price shocks in the USA are contained within the country (do not cause spill-overs to the EU and Oceania). On the basis of these results, the quantile dependence coefficients above are more likely to represent the probability that price shocks in the EU and in Oceania will stimulate shocks of similar magnitude in the USA rather than the opposite.

4. Conclusions

The present work has assessed the linkages between prices in the three major SMP producing regions (the EU, the USA and Oceania). The empirical

Table 4 Granger causality tests†

Causal pair of regions (A, B)	F -statistic	P -value
(USA, Oceania)	0.561	0.454
(Oceania, USA)	16.275	0
(USA, EU)	1.668	0.197
(EU, USA)	10.651	0.001
(Oceania, EU)	2.792	0.096
(EU, Oceania)	3.103	0.079

Note †Null hypothesis: price shocks from region A do not Granger-cause price shocks in region B.

investigation has relied on monthly price observations from 2001 to 2014 and on nonparametric kernel-based copulas which dispense with the need to specify functional forms for the relationships under study.

The empirical results suggest:

1. There has been a considerable degree of price co-movement (both an overall one as well as at the extremes of the respective joint distributions). Moreover, its strength has been time-varying (steadily rising) over the sample period. The rapid expansion of the global SMP market, the drive for trade liberalisation and the rise of the USA as a major exporter are possible explanations for these developments.
2. The spatial price linkages have been stronger for Oceania and the EU (the two producing regions which have traditionally competed on the same export markets). The USA, however, has been catching up with them as it has been gradually shifting the destination of its SMP exports from Mexico to China, South Asia, Middle East and North Africa.
3. Price co-movement with respect to the sign of price shocks has been symmetric in the sense that positive and negative shocks of the same absolute magnitude have been transmitted with the same intensity. Symmetry with respect to sign is an additional indicator of spatial market integration (Meyer and von Cramon-Taubadel 2004). A possible explanation for this finding is that SMP is sold in the international market as an undifferentiated commodity.
4. Larger (in absolute value) price shocks are likely to be transmitted from one region to another with higher intensity compared to smaller price shocks. This result is consistent with transaction costs literature and provides support to findings from econometric models pointing to price inertia around the medians and to the existence of inactivity bands.
5. Price shocks in Oceania and in the EU do have spill-over effects to the USA market. Price shocks in the USA, however, appeared to be contained within the country.

Trade intensification is expected to enhance price co-movement and integration of geographically separated markets. According to OECD-FAO Agricultural Outlook (2014), SMP trade will experience an annual growth rate of about 2 per cent from 2014 to 2023. This presupposes strong economic growth in the key SMP markets. Future developments are likely to be shaped by: (i) The impact of the removal, in April 2015, of the EU milk production quota on the supply of milk and of dairy commodities. In anticipation of the quota's abolition, more than \$1.2 billion have been invested in milk processing. Part of that amount has been already been directed to powder drying facilities for export markets (Dairy Australia, 2013). (ii) The efforts of major importers (China and Russian Federation) to build self-sufficiency in dairy commodities. (iii) The outcome of various multilateral and bilateral trade agreements currently under consideration. (iv)

The increased awareness with respect to environmental issues associated with dairy production (e.g. greenhouse-gas emissions, manure management). As a result, environmental constraints may become more binding for certain major exporters in the international dairy markets, notably New Zealand.

There are two potential avenues for future research. One may involve the application of kernel-based copulas to a larger number of geographically separated dairy producing regions. This, of course, presupposes the availability of price series with comparable length and quality. The second may involve comparison of the results from parametric copulas to those from kernel-based ones. Here, the statistical significance of the dependence coefficients at the extremes indicates that the Gaussian family and certain Archimedean families (e.g. the Frank copula) would be probably overly restrictive for modelling price co-movement in the international SMP market.

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Supporting Information

Additional Supporting Information may be found in the online version of this article:

Appendix S1. SMP Production and exports. Average 2011–2013

Appendix S2. Summary statistics: raw price shocks

Appendix S3. Estimates from the skewed- t ARMA-GARCH models*

Appendix S4. Natural logarithms of SMP prices

Appendix S5. The non-parametric test of time-invariant co-movement

Appendix S6. The non-parametric version of Hoeffding's ϕ

Data S1. Data & code fousekis—November 15.