AgEcon SEARCH
RESEARCH IN AGRICULTURAL & APPLIED ECONOMICS

The World's Largest Open Access Agricultural & Applied Economics Digital Library

This document is discoverable and free to researchers across the globe due to the work of AgEcon Search.

Help ensure our sustainability.

Give to AgEcon Search

AgEcon Search
http://ageconsearch.umn.edu
aesearch@umn.edu

*Papers downloaded from **AgEcon Search** may be used for non-commercial purposes and personal study only. No other use, including posting to another Internet site, is permitted without permission from the copyright owner (not AgEcon Search), or as allowed under the provisions of Fair Use, U.S. Copyright Act, Title 17 U.S.C.*

PRICE TRANSMISSION AND ASYMMETRIC ADJUSTMENT IN THE SPANISH DAIRY SECTOR

2002 AAEA-WAEA Annual Meeting

Teresa Serra*

Agricultural, Environmental and Development Economics Department, The Ohio State University
Economics Department, University of Vic

and

Barry K. Goodwin*

Agricultural, Environmental and Development Economics Department, The Ohio State University

January 31, 2002

Asymmetric threshold error correction models are used to analyze price relationships and patterns of transmission among farm and retail markets for a variety of dairy products in Spain. Our empirical analysis is conducted using both weekly and monthly price data. Our results suggest that the transmission of shocks is largely unidirectional, running from the farm to retail levels. Formal tests confirm asymmetries, though our results suggest that the effects of these asymmetries are modest. Implications for the organizational structure of Spanish dairy markets are offered.

Keywords: vertical price transmission, asymmetries, thresholds

JEL codes: Q110, Q130

---

*219 Agricultural Administration Bldg., Agricultural, Environmental and Development Economics Department, The Ohio State University, Columbus, OH 43210-1067. Tel: (614) 688 41 53. Fax: (614) 292 47 49. E-mail: serra.10@osu.edu.

Teresa Serra gratefully acknowledges research support from the Caja Madrid Foundation.

*229 Agricultural Administration Bldg., Agricultural, Environmental and Development Economics Department, The Ohio State University, Columbus, OH 43210-1067. Tel: (614) 688 41 38. Fax: (614) 292 47 49. E-mail: goodwin.112@osu.edu.

Copyright 2002 by Teresa Serra and Barry K. Goodwin. All rights reserved. Readers may make verbatim copies of this document for non-commercial purposes by any means, provided that this copyright notice appears on all such copies.

PRICE TRANSMISSION AND ASYMMETRIC ADJUSTMENT IN THE SPANISH DAIRY SECTOR

The Spanish milk sector has undergone important changes in recent years. The dairy industry has experienced many mergers and acquisitions, especially in the liquid milk sector, yielding higher levels of industry concentration. There has also been a reorientation of production; characterized by a tendency to increase the output of high value added goods to the detriment of less processed products. In spite of the active restructuring process, Spanish dairy industries continue to be very small in comparison to most retail firms. The retail sector has also realized significant changes, leading to higher concentration levels. The milk-producing sector has also experienced significant changes, with a decrease in the number of farms, an increase in concentration, a reduction in the dairy cow herd and an increase in production specialization and intensification. The European Union's Common Agricultural Policy (CAP) reforms over the last decade, intended to gradually reduce the price of surplus products by undertaking to compensate for farmer's lower incomes using direct aid packages, also constitute changes affecting the evolution of the Spanish dairy sector in recent years.

Price is the primary mechanism by which different levels of the market are linked. The analysis of vertical price transmission allows one to better understand the overall functioning of the market. The extent and speed with which shocks are transmitted between different levels of the marketing chain can have important implications for pricing practices and may reflect the level of competition in the market. In a competitive market with perfect information, price changes at one market level will usually cause changes in other levels. Market efficiency often suggests an equilibrium relationship between prices at different levels of the marketing chain. Some authors have hypothesized that the long-run relationship between prices may be asymmetric.[1] This may occur if middlemen in the marketing chain pass input price increases to customers more quickly and completely than input price reductions.

The study of vertical price transmission among various levels of the food market has recently gained special importance in the economics literature. The attention devoted to these analyses can be explained by two main factors. First, progressive concentration has been occurring both in the food industry and in the distribution sector. This concentration may modify the competitive positions of different economic agents participating in the market and may alter price transmission processes. A second point of relevance is the recent developments in time series econometrics. It is now recognized

[1] See, for example, Ward (1982), Bailey and Brorsen (1989), Kinnucan and Forker (1987) and Goodwin and Holt (1999).

that conventional methodologies that ignore nonstationarity may suffer from specification errors and inferential biases.

Recent research on price transmission has focused on the potential for asymmetries in the adjustment of prices at different levels of the market. Several theoretical and institutional reasons that may bring about asymmetries have been offered. First, Ward (1982) explains that those agents possessing perishable goods may not increase prices to avoid the risk of being left with spoiled product. Market power could be a second cause of asymmetry. Different costs of adjustment, depending on whether prices rise or fall, might be a third cause (Bailey and Brorsen 1989). Different price-elasticities at different levels of the marketing chain may be a fourth reason. Finally, public intervention to support producer prices could also cause asymmetry (Kinnucan and Forker 1987).

A significant set of analyses addressing the asymmetry question has involved the use of variations of the econometric specification introduced by Wolffram (1971) and refined by Houck (1977) and Ward (1982).[2] Although a generalization of the results of these analyses is somewhat difficult to make, most research has detected asymmetries in price adjustments at different market levels, although the extent of asymmetry is generally small. Additionally, most existing research has found that price changes tend to flow from the farm to wholesale and retail markets. Farm prices rarely respond to wholesale or retail price shocks.

Specifications that use some variation of the Wolffram (1971) method have been criticized because they ignore the time-series properties of the data. In particular, many analyses have not considered the problems associated with nonstationary data. To adequately study asymmetry of price transmission, Cramon-Taubadel (1998) proposed a modification of the standard Wolffram specification to allow for an error correction term. He found evidence of asymmetries in price adjustment in German producer and wholesale hog markets. More recently, Goodwin and Holt (1999) proposed the use of threshold vector error correction models (TVECM) to take into account the potential for nonlinear and threshold-type adjustments in error correction models. The TVECM model is a multivariate version of threshold autoregressive (TAR) models. TVECM models allow one to investigate the adjustment process of individual prices and provide more information about short-run price dynamics.[3]

Balke and Fomby (1997) introduced threshold cointegration and error correction models. They suggested a grid search procedure whereby threshold parameters are chosen by minimizing a sum of

[2] See, for example, Heien (1980), Kinnucan and Forker (1987) or Bailey and Brorsen (1989).

[3] TVECM models have also been used by Lo and Zivot (1999), Goodwin and Harper (2000), Goodwin and Piggott (2001), and Goodwin, Grennes and Craig (2001).

squared errors (SSE) criterion. In the context of multivariate models, such an approach may be less preferred to a criterion that recognizes the potential non-independence of residuals across equations.[4]

Our analysis evaluates vertical price transmission patterns in Spanish milk markets. We concentrate on the relationships between producer and retail prices. We use three-regime threshold vector error correction models that allow asymmetric price adjustments and reflect vertical price dynamics. We introduce an econometric procedure to estimate the threshold parameters that accounts for the relationship between markets at different levels of the marketing chain. Specifically, to determine the threshold parameters, we use a grid search that minimizes the logarithm of the determinant of the Σ matrix, which is analogous to maximizing a conventional Likelihood Ratio (LR) Chow test. We compare our results to those derived from the most common methodology based on the minimization of the trace of the Σ matrix (i.e., by minimizing system SSE).

Econometric methods

Tong (1978) originally introduced nonlinear threshold time series models. Tsay (1989) developed a method to test for threshold effects in autoregressive models and to model threshold autoregressive processes. Balke and Fomby (1997) extended the threshold autoregressive models to a cointegration framework.

Consider a standard linear cointegration relationship between a pair of prices (P_{it} and P_{jt}).

$$P_{it} - aP_{jt} = n_t$$

where $n_t = f n_{t-1} + u_t$ represents the residual of the equilibrium relationship (i.e. a deviation from equilibrium). Cointegration between P_{it} and P_{jt} requires n_t to be stationary, implying $|f|<1$. Balke and Fomby (1997) extended this analysis to the case where n_t follows a threshold autoregression. If a three-regime threshold autoregressive model is used, the behavior of n_t can be modeled as:

$$n_t = f n_{t-1} + u_t$$

[4] Put a different way, when estimating the threshold parameters, previous analyses have not exploited the information contained in the variance-covariance (Σ) matrix of the residuals of the TVECM, to take into account the possible relationships among the markets being analyzed. Assuming cross-sectional independence among the residuals of the model, these analyses have used a grid search aimed at minimizing the trace of the Σ matrix to estimate the threshold parameters. Hence, the influence of the residual covariances has not been considered.

$$\begin{cases} f^{(1)} & if \quad -\infty \langle n_{t-d} \leq c_1 \\ f^{(2)} & if \quad c_1 \langle n_{t-d} \leq c_2 \\ f^{(3)} & if \quad c_2 \langle n_{t-d} \leq +\infty \end{cases}$$

where c_1 and c_2 represent the thresholds that delineate the different regimes and n_{t-d} represents the variable relevant to the threshold behavior. As in most empirical applications, we assume that d is equal to 1.

The vector error correction representation of the threshold model is given by:

$$\Delta P_t = \begin{cases} \sum_{i=1}^{l} B_i^{(1)} \Delta P_{t-i} + p^{(1)} n_{t-1} + e_t^{(1)} & if \quad -\infty \langle n_{t-d} \leq c_1 \\ \sum_{i=1}^{l} B_i^{(2)} \Delta P_{t-i} + p^{(2)} n_{t-1} + e_t^{(2)} & if \quad c_1 \langle n_{t-d} \leq c_2 \\ \sum_{i=1}^{l} B_i^{(3)} \Delta P_{t-i} + p^{(3)} n_{t-1} + e_t^{(3)} & if \quad c_3 \langle n_{t-d} \leq +\infty \end{cases}$$

where P_t is the vector of prices being analyzed (P_{it} and P_{jt}). The TVECM can be compactly expressed as:

$$\Delta P_t = \begin{cases} B^{(1)'} x_{t-1} + e_t^{(1)} & if \quad -\infty \langle n_{t-d} \leq c_1 \\ B^{(2)'} x_{t-1} + e_t^{(2)} & if \quad c_1 \langle n_{t-d} \leq c_2 \\ B^{(3)'} x_{t-1} + e_t^{(3)} & if \quad c_2 \langle n_{t-d} \leq +\infty \end{cases}$$

where:

$$x_{t-1} = \begin{bmatrix} \Delta P_{t-1} \\ \Delta P_{t-2} \\ . \\ . \\ . \\ \Delta P_{t-l} \\ n_{t-1} \end{bmatrix}$$

This may also be written as:

$$\Delta P_t = b^{(1)'} x_{t-1} d_{1t}(c_1, c_2, d) + b^{(2)'} x_{t-1} d_{2t}(c_1, c_2, d) + b^{(3)'} x_{t-1} d_{3t}(c_1, c_2, d) + e_t$$

where:

$$d_{1t}(c_1, c_2, d) = 1(-\infty < v_{t-d} \le c_1)$$
$$d_{2t}(c_1, c_2, d) = 1(c_1 < v_{t-d} \le c_2)$$
$$d_{3t}(c_1, c_2, d) = 1(c_2 < v_{t-d} \le +\infty)$$

Our specific estimation strategy can be summarized as follows. First, in order to determine whether the price series are stationary, standard Dickey-Fuller and Kwiatkowski et al. (1992) unit root tests are used.[5] Second, we test for cointegration among the prices studied using the Johansen cointegration test. The test is carried out for each pair of raw-milk and manufactured dairy product prices. We then follow the general approach of Engle and Granger and utilize ordinary least squares estimates of the cointegration relationships among the pairs of prices. The lagged residuals derived from these relationships are used to define the error correction terms. The next step consists of determining whether the dynamics of the long-run relationships among prices are linear or whether they exhibit threshold-type nonlinearities. We use Tsay's (1989) nonparametric test.

In order to estimate the parameters of the multivariate TVECMs we use sequential conditional iterated SUR in two steps. In the first step, a two-dimensional grid search is carried out to estimate the threshold parameters (c_1 and c_2). The thresholds are searched over 1% and 99% of the fractiles of the negative and positive lagged error correction terms. The search is restricted to ensure an adequate number of observations for estimating the parameters in each regime. Two alternative grid search techniques are utilized. The first minimizes the log determinant of the variance-covariance (Σ) matrix of the residuals of the TVECMs, which is analogous to maximizing a standard LR test. This criteria differs from the one used in other analyses in that it does not assume cross-equation independence between the residuals.[6] In the former criterion, we account for the relationship between markets at different levels of the marketing chain. We compare results obtained from both procedures.

The vectors of parameters $B^{(1)'}$, $B^{(2)'}$ and $B^{(3)'}$ are estimated by iterated SUR method giving:

[5] The Kwiatkowski et al. (1992) test for stationarity is also applied to overcome the inconvenience that the standard test entails: unless there is strong evidence against the null hypothesis, it is accepted.

[6] See for example Obstfeld and Taylor (1997), Goodwin and Holt (1999), Lo and Zivot (1999) and Goodwin and Piggott (2001).

$$S(c_1, c_2, d) = \ln\left|\hat{\Sigma}(c_1, c_2, d)\right|$$

if the first criteria is used, or

$$S(c_1, c_2, d) = trace\left(\hat{\Sigma}(c_1, c_2, d)\right)$$

if the second criteria is used, where $\hat{\Sigma}(c_1, c_2, d)$ is a multivariate iterated SUR estimate of Σ=var(e_t) conditional on (c_1,c_2,d) and d is assumed to be equal to 1. In the second step, the estimates of c_1 and c_2 are obtained as:

$$(\hat{c}_1, \hat{c}_2, 1) = \arg \min_{c_1, c_2} S(c_1, c_2, 1)$$

The final estimates of the parameters are given by $\hat{B}^{(i)'} = \hat{B}^{(i)'}(\hat{c}_1, \hat{c}_2, 1)$ and the estimations of the residual covariance matrix by $\hat{\Sigma} = \hat{\Sigma}(\hat{c}_1, \hat{c}_2, 1)$.

Finally, we test for the significance of the differences in parameters across relative regimes. We use the sup-LR statistic to test for a linear VECM against the alternative of a TVECM, which is an extension of the Hansen's approach to test for the statistical significance of the threshold effects in univariate TAR models to a multivariate TVECM. The model under the null is $\Delta P_t = B' x_{t-1} + e_t$, while the model under the alternative can be written as:

$$\Delta P_t = b^{(1)'} x_{t-1} d_{1t}(c_1, c_2, d) + b^{(2)'} x_{t-1} d_{2t}(c_1, c_2, d) + b^{(3)'} x_{t-1} d_{3t}(c_1, c_2, d) + e_t.$$

The sup-LR statistic can be computed in the following way:

$$LR = T\left(\ln\left|\hat{\Sigma}\right| - \ln\left|\hat{\Sigma}(\hat{c}_1, \hat{c}_2, 1)\right|\right),$$

where $\hat{\Sigma}$ is the covariance matrix of the residuals of the VECM, $\hat{\Sigma}(\hat{c}_1, \hat{c}_2, 1)$ represents the covariance matrix of the residuals of the TVECM, and T is the number of observations. The sup-LR statistic has a non-standard distribution because the threshold parameters are not identified under the null hypothesis. To determine the p-value of the sup-LR statistic, we run 100 simulations for each model whereby the dependent variables (ΔP_t) are replaced by iid N(0,1) draws (see Hansen (1997) for a detailed discussion of this approach). The proportion of simulations under the null for which the simulated LR statistic exceeds the observed LR statistic gives the asymptotic p-value of the sup-LR test.

Empirical application

Our empirical analysis utilizes dairy prices in Spain, observed from the last week of June 1994 to the last week of December 2000. The period of study follows the 1992 MacSharry reforms, which moved the CAP toward income support through compensatory direct payments, allowing a reduction in guaranteed prices[7]. Weekly retail prices were taken from *Alimentación Precios Venta al Público* database, provided by the Spanish Ministry of Economy[8]. They include the following manufactured products: pasteurized liquid milk, sterilized liquid milk, condensed milk, powdered milk, continuation milk, fresh cheese, blended cheese, manchego cheese, dutch cheese, cheese in portions, emmenthal cheese, yogurt, cream-caramel and butter. Producer prices for raw milk were also collected on a weekly basis from the database *Precios Testigo Nacionales* of the Spanish Ministry of Agriculture, though an important distinction regarding these prices must be noted.[9] Because of policy and marketing arrangements, producer prices typically do not change on a week-to-week basis. Rather, the prices are constant throughout the month with adjustments being made at the beginning of each month.[10] Although our sources report prices on a weekly basis, the reported prices are constant across a month. In this light, we consider two alternative phases of the analysis. In the first, we utilize the weekly reported prices. In the second, monthly average prices are used.

Standard unit-root tests confirm the presence of a unit root in all weekly price series except for powdered and continuation milk and blended, manchego and portions cheeses. The Kwiatkowski tests support the presence of a unit root in every time series. When standard unit-root tests are applied to monthly price frequency and monthly dummies to allow seasonality are introduced, we obtain evidence of unit roots in all price series except for condensed milk and blended cheese. Johansen cointegration tests

[7] In 1999, another reform (Agenda 2000) was launched. Agenda 2000, based on the principles established in 1992, insisted on a more competitive and market-oriented agricultural sector by further reducing prices for some commodities and compensating producers with direct payments. Due to budget constraints, the policy measures approved for the dairy sector will not be applied until 2005.
[8] Retail prices are average national prices paid by the consumers for dairy products. We have taken the "frequent" prices, which means that they correspond to an average quality. These prices include the value added tax, which has been removed.
[9] Raw milk prices correspond to a standard quality milk.
[10] This point was confirmed by talks with dairy industries and cooperatives.

(table 1) indicate[11] the existence of a single cointegration relationship among all pairs of weekly prices except for raw milk-blended cheese; raw milk-butter; raw milk-emmenthal cheese and raw milk-fresh cheese.[12] When the analysis is repeated for monthly prices, the results are identical. Lag orders for cointegration tests were selected using SC and HQ information criteria. The deterministic components introduced in the vector error correction model to carry out the Johansen test were chosen according to information criteria and the Likelihood Ratio test.

Though Balke and Fomby (1997) showed that standard methods for evaluating unit roots and cointegration may work reasonably well when prices are linked by a threshold cointegration relationship, they, along with Enders and Granger (1998), have also shown that standard cointegration tests may lack power in the presence of asymmetric adjustment. We compute the OLS estimates of the error correction terms for all pairs of variables using the price of raw milk as the normalization variable.[13]

Tsay's test is conducted using the error correction terms derived from the OLS cointegration regression. When weekly data are used, Tsay's test supports nonlinearity at the 10% significance level in all models except for the raw milk-continuation milk model (see table 2). In the monthly price analysis, nonlinearity is not implied for any model (see table 3). The thresholds derived from the two dimensional grid searches and the sup-LR statistics are presented in tables 2 for weekly data and 3 for monthly price frequency. When weekly data are used, our results suggest that threshold effects are statistically significant at the 10% level for the following models: raw milk-blended cheese, raw milk-condensed milk, raw milk-continuation milk, raw milk-manchego cheese, raw milk-portions cheese and raw milk-powdered milk. For the rest of the models the sup-LR test rejects asymmetries in the process of price adjustment. Hence, in contrast to Ward's (1982) argument, no asymmetries are apparent in the transmission of shocks in highly perishable dairy product prices. Likewise, the results do not suggest a relationship between asymmetry and concentration in the dairy industry[14]. This result is not surprising because, as mentioned above, dairy industries have little market power relative to the big retail chains. Additionally, it should be noted that a conspicuous part of the milk collected by the industry in Spain is earmarked for the production of liquid milk. While in the E.U. a quarter of the milk collected in 1996 was

[11] If the λ-max and λ-trace tests lead to different results, the first statistic is used. Johansen and Juselius (1990) suggest that the λ-trace test may lack power relative to the maximal eigenvalue test, indicating that the trace test is more likely to support the presence of cointegrating vectors.

[12] Empirical evidence suggesting that the prices for milk and butter do not have a linear long-run relationship may suggest that the intervention prices for butter (defined in the framework of the EU's Common Agricultural Policy) may not exert strong influence on the prices received by milk producers in Spain. A weak relationship between market and intervention prices is confirmed by Foro Agrario (2000) for the whole E.U.

[13] Results are available from the authors upon request.

[14] According to Alimarket data, the concentration level in the cheese branch is amongst the lowest in the dairy industry. In spite of this, we have detected asymmetries for blended, manchego and portions cheeses. On the other hand, even though the yogurt and dairy dessert retail outlets in Spain are highly concentrated, no threshold effects have been detected for these products.

intended for consumption in the form of liquid milk, in Spain this use accounted for more than 60% (Ministère de l'Agriculture et de la Pêche 2000, 162). The Spanish liquid milk industry is characterized by low value added to factor costs, a reduced market price of the manufactured product, a scarcity of raw milk and a small degree of market power relative to the big food retail chains. These characteristics may contribute to tightening the relationship between the evolution of the retail and farm prices, reducing the possibility of an asymmetric price adjustment.

When monthly data are used, the results are very similar. However, the results differ for the raw milk-cream caramel model, for which the sup-LR test suggests statistically significant threshold effects. The sup-LR test results also support a three-regime TVECM for raw milk-fresh cheese and raw milk-pasteurized milk if the thresholds are chosen by minimizing the log determinant of the Σ matrix. On the other hand, the sup-LR statistic indicates that the thresholds are not significant for the raw milk-manchego cheese model. Hence, when monthly prices are used, the empirical results provide more support to Ward's hypothesis.

An important finding is that the two grid search techniques that have been used do not lead to significant differences in the results when weekly data are used. If the analysis is performed with monthly prices, a couple of relevant differences arise. The differences are present in the raw milk-fresh cheese and the raw milk-pasteurized milk models. While the sup-LR test supports asymmetries if the models are estimated minimizing the natural log of the determinant of the Σ matrix, asymmetries are rejected if the more common methodology (based on the minimization of the trace of the Σ matrix is used). In order to better interpret the dynamic relationships among prices, impulse response functions are considered for those models with significant threshold effects. In threshold models, responses to a shock depend on the history of the series, as well as on the size and sign of the shock. As a consequence, many different impulse response functions can be computed. We chose a single observation (the last observation of our data) to evaluate the responses to one standard deviation positive and negative shocks to the price of raw milk and to the price of each manufactured dairy products. We adopt Koop, Pesaran and Potter's (1996) proposal, which defines responses (I_{t+k}) on the basis of the observed data (z_t, z_{t-1}, …) and a shock (v) as:

$$I_{t+k}(v, Z_t, Z_{t-1}, \ldots) = E\left[Z_{t+k} \middle| Z_t = z_t + v, Z_{t-1} = z_{t-1}, \ldots\right] - E\left[Z_{t+k} \middle| Z_t = z_t, Z_{t-1} = z_{t-1}, \ldots\right]$$

Because some of the prices are not stationary, we may find transitory as well as permanent responses. Specifically, for those pairs of variables with a nonstationary error correction term, shocks may provoke a permanent alteration of the variables' time path.

The responses to one standard deviation positive and negative shocks to the monthly price of the dairy manufactured products are illustrated in figure 1[15]. The responses to one standard deviation positive and negative shocks to the monthly price of raw milk are illustrated in figure 2[16]. An implication of the impulse responses is that, though parametric differences across relative regimes are statistically significant, price adjustments appear to be reasonably symmetric for both weekly and monthly data. Hence, in contrast to Kinnucan and Forker's (1987) argument, public regulation of farm prices does not appear to lead to asymmetric vertical price transmission in the Spanish dairy sector. Our results are consistent with the similar findings of Goodwin and Holt (1999) for U.S. beef markets and Boyd and Brorsen (1988) for U.S. pork markets. However our finding of limited asymmetries is in contrast with the findings of Cramon-Taubadel (1998) for German pork markets, Hahn (1990) for U.S. beef markets and Kinnucan and Forker (1987) for U.S. dairy markets.

A second implication of the impulse responses is that, when weekly data are used, there is little feedback to farm prices from shocks at the retail level. On the other hand, farm market price shocks usually elicit responses at the retail level. The modest response of farm prices to retail price shocks may be in part due to organizational characteristics of the Spanish dairy sector. The Spanish milk sector is characterized by an overall lack of marketing contracts and a relative scarcity of farmer cooperatives. While only 30% of milk in Spain is marketed through cooperatives, this proportion reaches 68% in the E.U. (Foro Agrario 2000, 216). As a result, the prices for a high proportion of the milk collected are established by direct negotiation between the dairy industry and the farmer. This implies that the farm price is mainly determined by the industry, due to the little market power of the farmers relative to the dairy industry (Foro Agrario 2000, 317). This result may also reflect the fact that weekly farm prices are fixed within the month, as is noted above. When monthly data are used, farm prices appear to be more elastic in response to shocks in retail prices.

Price shocks seem to provoke permanent adjustments in most of the models, reflecting the nonstationary nature of the price data. Retail price shocks bring about a permanent adjustment to retail prices and, in some cases, to farm prices. Permanent adjustments also characterize most responses of farm and retail prices to farm shock prices. The responses of retail markets to farm price shocks appear, in most of the cases, to be somewhat damped during the initial period after the shock as one moves up the marketing chain (farm prices exhibit, in absolute values, the largest response, followed by retail price adjustments). After the first period, retail prices exhibit, in absolute values, the largest responses. This

[15] The TVECM that minimizes the natural log of the Σ matrix has been used to compute the impulse-response functions.

[16] The results derived from the monthly and weekly analyses are very similar. To conserve space, we have only presented the impulse-responses for monthly data. Weekly results are available from the authors upon request.

may indicate a slow adjustment of retail prices to farm price shocks, a result which is in accordance with Kinnucan and Forker (1987).

Some price shocks do not have the sign that one would expect. For weekly data, the unexpected signs are present in the raw milk-blended cheese, raw milk-continuation milk, raw milk-manchego cheese and raw milk-portions cheese models. Positive (negative) milk price shocks generate negative (positive) manufactured price shocks. This behavior, though somewhat difficult to interpret, may reflect the role of milk quotas and industry concentration. In particular, various component industries may compete to increase both their access to milk quota (i.e. to raw material) and their retail market share. This behavior could explain the increase in farm prices while the retail prices are reduced. The price of raw milk also responds with an unexpected sign to shocks to cheese in portions price. Similar unexpected signs are also found when the analysis uses monthly data frequency. It should be pointed out that, in some cases, the impulses suggest unstable responses. These anomalies may reflect a lack of cointegration between the prices analyzed.

Concluding remarks

We analyze price relationships and patterns of transmission among farm and retail markets for dairy products in Spain, using both weekly and monthly price data. Our analysis focuses on the time series properties of the prices. We estimate threshold vector error correction models with three regimes, which allow asymmetries in price transmission and recognizes the nonstationary nature of the price series.

Our results largely confirm the findings of research for other markets and commodities. In particular, we find that, when weekly data are used, the transmission of shocks appears to be largely unidirectional: retail prices adjust to farm level shocks to the raw milk price, but the milk price only modestly responds to retail market shocks. The weak response of farm prices to retail price shocks may be partly explained by the organizational characteristics of the Spanish milk sector. The lack of an organized contracting system and a scarcity of dairy farmer cooperatives may limit the market power of farmers relative to the dairy industry, as well as their capacity to negotiate prices. With monthly data, farm prices appear to be more elastic to shocks to retail prices, indicating that milk prices require a considerable long period to adjust. This result likely reflects the fact that farm prices are not adjusted on a weekly basis but rather change each month.

When weekly data are used, formal testing suggests the presence of asymmetries in vertical price transmission patterns for some manufactured dairy products with a relatively long shelf life. Asymmetries are not present in the price transmission of highly perishable dairy products. The lack of asymmetric relationships may reflect characteristics of liquid milk production in Spain. Liquid milk is a product characterized by low value added. This may tighten the relationship between retail and farm prices and could prevent asymmetric price adjustments. We have observed that the asymmetries do not appear to be related to the level of concentration of the dairy industry, which is expected in light of the limited power of the dairy industry to negotiate prices with large retail chains. When monthly data are used, the empirical results provide more support to Ward's suggestion that asymmetries may be found in the price transmission of highly perishable products.

Though formal tests confirm asymmetry for some products, an evaluation of the nonlinear impulse-response functions suggests that these differences are modest and may be economically insignificant. Hence, public regulation of farm prices does not apparently lead to asymmetric vertical price transmission in the Spanish dairy sector.

References

Bailey, D. and B.W. Brorsen. “Price asymmetry in spatial fed cattle markets.” *Western Journal of Agricultural Economics* 14 (2) (1989): 246-252.

Balke, N.S. and T.B. Fomby. “Threshold cointegration.” *International Economic Review*, vol. 38, no 3 (1997): 627-645.

Boyd, M.S. and B.W. Brorsen. “Price Asymmetry in the U.S. Pork Marketing Channel.” *North Central Journal of Agricultural Economics* 10 (January 1998): 103-110.

Von Cramon-Taubadel, S. “Estimating price transmission with the error correction representation: an application to the German pork market.” *European Review of Agricultural Economics* 25 (1998): 1-18.

Enders, W. and C.W.J. Granger. “Unit root tests and asymmetric adjustment with an example using the term structure of interest rates.” *Journal of Business and Economic Statistics* 16(1998): 304-311.

Engle, R.F. and C.W.J. Granger. “Cointegration and error correction: representation, estimation and testing.” *Econometrica* 55 (1987): 251-276.

Foro Agrario. *La Reforma de la PAC de la Agenda 200 y la Agricultura Española.* Ediciones Mundi-Prensa, 2000.

Goodwin, B. K. and M. T. Holt. ”Asymmetric Adjustment and Price Transmission in the U.S. Beef Sector.” *American Journal of Agricultural Economics* (principal paper) 81(1999):630-37.

Goodwin, B. K. and D. Harper. ”Price Transmission, Threshold Behavior, and Asymmetric Adjustment in the U.S. Pork Sector.” *Journal of Agricultural and Applied Economics* 32(December 2000).

Goodwin, B. K. and N. Piggott. "Spatial Market Integration in the Presence of Threshold Effects." *American Journal of Agricultural Economics* May 2001.

Goodwin, B. K., T. Grennes, Lee Craig. ”Mechanical Refrigeration and the Integration of Perishable Commodity Markets.” accepted for publication in Explorations in Economic History, 2002.

Hahn, W.F. “Price Transmission Asymmetry in Pork and Beef Markets”. *The Journal of Agricultural Economic Research* 42(1990): 21-30.

Houck, P.J. “An approach to specifying and estimating non reversible functions.” *American Journal of Agricultural Economics* 59 (1977): 570-572.

Hansen, B.E. “Inference in TAR models.” *Studies in Nonlinear Dynamics Econometrics* 2(1) (April 1997): 1-14.

Johansen, S. “Statistical analysis of cointegrating vectors.” *Journal of Economic Dynamics and Control* 12 (1988): 231-254.

Kinnucan, H.W. and O.D. Forker. “Asymmetry in farm-retail price transmission for major dairy products.” *American Journal of Agricultural Economics* 69 (1987): 285-292.

Koop, G., M.H. Pesaran, and S.M. Potter. "Impulse response analysis in nonlinear multivariate models." *Journal of Econometrics* 74 (1996): 119-146.

Kwiatkowski, D. et al. "Testing the null hypothesis of stationarity against the alternative of a unit root: how sure are we that economic time series have a unit root?." *Journal of Econometrics* (Oct. 1992): 159-178.

Lo, M.C. and E. Zivot. "Threshold cointegration and nonlinear adjustment to the law of one price." *Macroeconomic Dynamics*, forthcoming.

Ministère de l'Agriculture et de la Pêche. *Agriculture in the European Union*. Agreste, 2000.

Tong, H. *On a threshold model in pattern recognition and signal processing*. C.Chen, ed. Amsterdam: Sijhoff and Noonhoff, 1978.

Tsay, R. "Testing and modelling threshold autoregressive processes." *Journal of the American Statistical Association* 84 (1989): 231-240.

Ward, RW. "Asymmetry in retail, wholesale, and shipping point pricing for fresh vegetables." *American Journal of Agricultural Economics* 62 (1982): 205-212.

Wolffram, R. "Positive measures of aggregate supply elasticities: some new approaches-some critical notes." *American Journal of Agricultural Economics* 53 (1971): 356-359.

Table 1. The Johansen cointegration tests

Model	Cointegration tests weekly frequency				Cointegration tests weekly frequency			
	λ max (sig. value 90%) r = 0	λ max (sig. value 90%) r = 1	λ trace (sig. value 90%) r = 0	λ trace (sin. value 90%) r = 1	λ max (sig. value 90%) r = 0	λ max (sig. value 90%) r = 1	λ trace (sig. value 90%) r = 0	λ trace (sin. value 90%) r = 1
Raw milk-Blended cheese	29.93 (10.29)	8.21 (7.50)	38.14 (17.79)	8.21 (7.50)	30.18 (10.29)	7.96 (7.50)	38.14 (17.79)	7.96 (7.50)
Raw milk-Butter	17.30 (10.29)	8.07 (7.50)	25.37 (17.79)	8.07 (7.50)	18.63 (10.29)	7.88 (7.50)	26.51 (17.79)	7.88 (7.50)
Raw milk-Condensed milk	23.78 (12.39)	8.27 (10.56)	32.05 (22.95)	8.27 (10.56)	22.75 (12.39)	8.45 (10.56)	31.20 (22.95)	8.45 (10.56)
Raw milk-Continuation milk	45.67 (10.29)	4.74 (7.50)	50.41 (17.79)	4.74 (7.50)	55.01 (10.29)	3.85 (7.50)	58.86 (17.79)	3.85 (7.50)
Raw milk-Creamcaramel	13.45 (12.39)	6.50 (10.56)	19.96 (22.95)	6.50 (10.56)	16.23 (12.39)	6.80 (10.56)	23.03 (22.95)	6.80 (10.56)
Raw milk-Dutch cheese	25.53 (12.39)	4.10 (10.56)	29.63 (22.95)	4.10 (10.56)	22.29 (12.39)	5.15 (10.56)	27.44 (22.95)	5.15 (10.56)
Raw milk-Emmenthal cheese	13.97 (10.29)	7.53 (7.50)	21.50 (17.79)	7.53 (7.50)	15.47 (10.29)	8.92 (7.50)	24.39 (17.79)	8.92 (7.50)
Raw milk-Fresh cheese	17.54 (12.39)	14.73 (10.56)	32.26 (22.95)	14.73 (10.56)	19.15 (12.39)	16.29 (10.56)	35.44 (22.95)	16.29 (10.56)
Raw milk-Manchego cheese	23.72 (12.39)	7.72 (10.56)	31.44 (22.95)	7.72 (10.56)	20.20 (12.39)	9.11 (10.56)	29.32 (22.95)	9.11 (10.56)
Raw milk-Pasteurized milk	18.33 (12.39)	2.66 (10.56)	21.00 (22.95)	2.66 (10.56)	14.14 (10.29)	7.50 (7.50)	21.63 (17.79)	7.50 (7.50)
Raw milk-Cheese in portions	29.45 (12.39)	4.70 (10.56)	34.15 (22.95)	4.70 (10.56)	27.19 (12.39)	4.46 (10.56)	31.65 (22.95)	4.46 (10.56)
Raw milk-Powdered milk	18.86 (12.39)	8.03 (10.56)	26.90 (22.95)	8.03 (10.56)	17.54 (12.39)	9.53 (10.56)	27.07 (22.95)	9.53 (10.56)
Raw milk-Sterilized milk	15.65 (10.29)	2.94 (7.50)	18.59 (17.79)	2.94 (7.50)	15.17 (10.29)	2.44 (7.50)	17.61 (17.79)	2.44 (7.50)
Raw milk-Yogurt	14.47 (12.39)	7.35 (10.56)	21.82 (22.95)	7.35 (10.56)	17.10 (12.39)	8.01 (10.56)	25.12 (22.95)	8.01 (10.56)

where:
λ MAX = Maximum eigenvalue test statistic;
λ TRACE = Trace test statistic;
r = number of cointegrating vectors being tested under the null hypothesis.

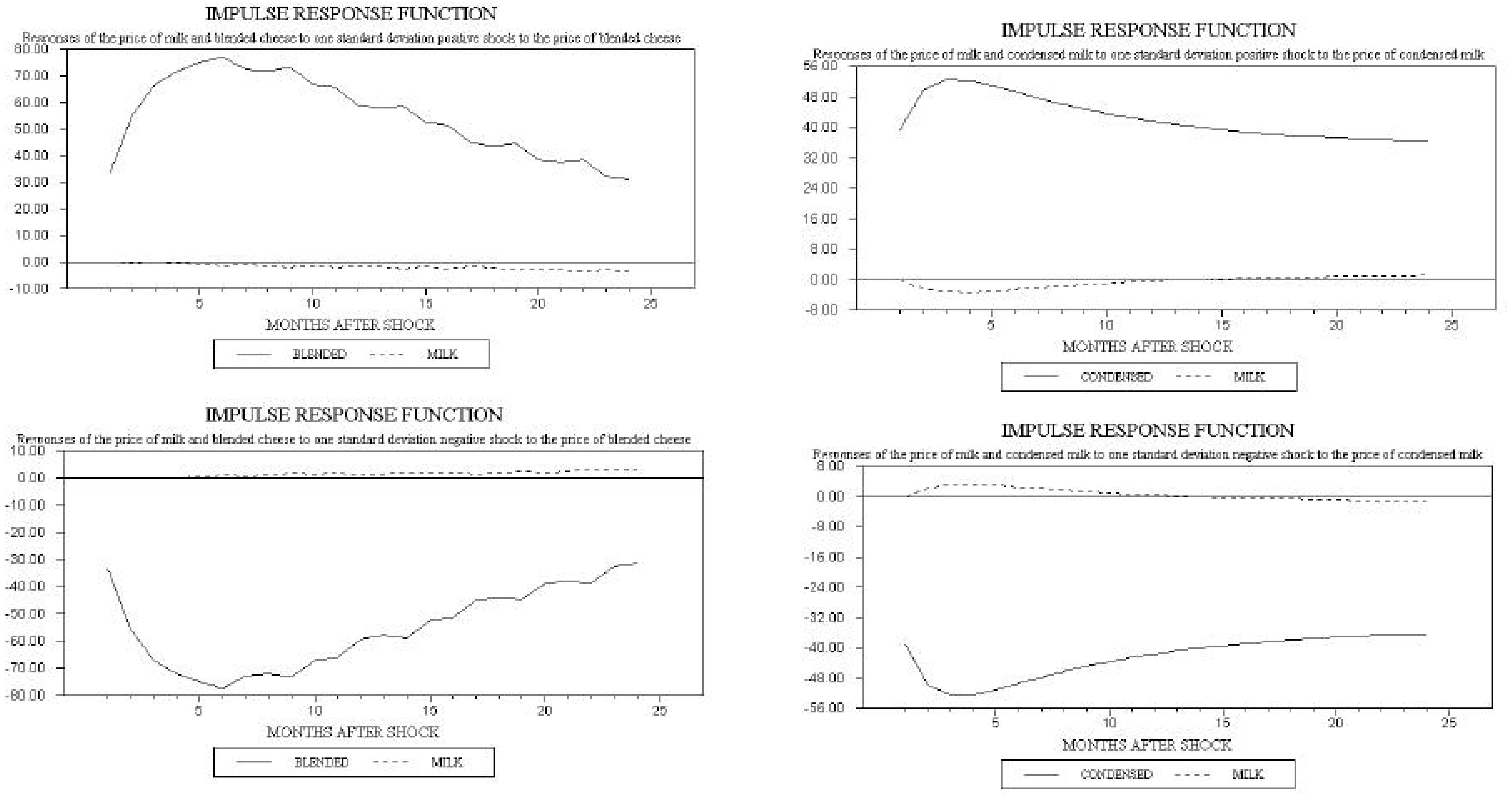
Table 2. Tsay's test, thresholds and the sup-LR test. Weekly frequency

Variables	Tsay's test (p-value)	Minimum ln|Σ|			Minimum trace(Σ)		
		C1: Negative threshold	C2: Positive threshold	Sup-LR test (p-value)	C1: Negative threshold	C2: Positive threshold	Sup-LR test (p-value)
Raw milk-Blended cheese	4.16592 (0.04202596)	-2.34095	1.62811	56.23925 (0.00000)	-2.35377	1.62811	56.23925 (0.00000)
Raw milk-Butter	3.37955 (0.06689279)	-0.03917	0.20428	31.77872 (0.60000)	-0.03917	0.20428	31.77872 (0.78000)
Raw milk-Condensed milk	4.51923 (0.03424479)	-0.18949	0.11953	58.02873 (0.02000)	-0.18949	0.11953	58.02873 (0.00000)
Raw milk-Continuation milk	0.83613 (0.36116062)	-0.24159	0.08563	135.47016 (0.00000)	-0.24159	0.08563	135.47016 (0.00000)
Raw milk-Creamcaramel	5.30740 (0.02184488)	-0.09037	0.17176	21.31526 (1.00000)	-0.72898	0.99290	19.61299 (1.00000)
Raw milk-Dutch cheese	7.46014 (0.00664105)	-1.15275	1.10033	36.04987 (0.47000)	-0.02112	1.96910	
Raw milk-Emmenthal cheese	4.67431 (0.03132250)	-0.34217	1.35144	18.97571 (1.00000)	-0.04831	0.46693	
Raw milk-Fresh cheese	6.30026 (0.01254156)	-1.00968	0.34257	34.26273 (0.57000)	-0.01618	1.41072	
Raw milk-Manchego cheese	8.34363 (0.00412151)	-1.62196	0.48595	55.05801 (0.02000)	-1.66291	0.48595	
Raw milk-Pasteurized milk	6.93376 (0.00884928)	-1.94840	0.11617	34.12964 (0.63000)	-0.91622	1.02559	
Raw milk-Cheese in portions	2.73864 (0.09888177)	-0.22173	0.49754	45.06852 (0.09000)	-0.22173	0.49754	45.06852 (0.11000)
Raw milk-Powdered milk	6.98988 (0.00858156)	-0.24999	1.18824	46.99935 (0.04000)	-0.24999	1.17721	46.99935 (0.04000)
Raw milk-Sterilized milk	5.24777 (0.02259362)	-1.19304	0.27564	23.29508 (1.00000)	-2.21721	0.24164	22.49824 (1.00000)
Raw milk-Yogurt	7.40394 (0.00684698)	-0.28832	1.31982	39.34844 (0.28000)	-0.28832	1.31982	39.34844 (0.23000)

Table 3. Tsay's test, thresholds and the sup-LR test. Monthly frequency

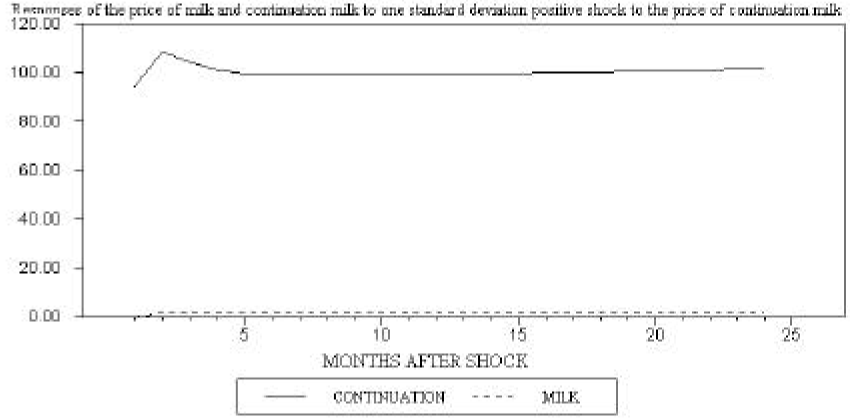
Variables	Tsay's test (p-value)	Minimum ln|Σ|			Minimum trace(Σ)		
		C1: Negative threshold	C2: Positive threshold	Sup-LR test (p-value)	C1: Negative threshold	C2: Positive threshold	Sup-LR test (p-value)
Raw milk-Blended cheese	0.84624 (0.36056708)	-1.14921	1.41948	40.09329 (0.00000)	-0.77494	1.54807	35.55594 (0.02000)
Raw milk-Butter	0.93321 (0.33713377)	-1.04811	0.82064	23.02497 (0.34000)	-1.08778	0.90279	21.16325 (0.48000)
Raw milk-Condensed milk	0.86902 (0.35421823)	-0.71381	0.61240	43.61249 (0.00000)	-0.71381	0.61240	43.61249 (0.00000)
Raw milk-Continuation milk	0.05845 (0.80961600)	-0.67959	0.38763	53.26109 (0.00000)	-0.76361	0.09217	52.00571 (0.00000)
Raw milk-Creamcaramel	0.69850 (0.40594238)	-0.78398	1.06550	32.92471 (0.09000)	-0.78398	1.06550	32.92471 (0.06000)
Raw milk-Dutch cheese	0.20474 (0.65223010)	-0.74362	0.28103	23.27535 (0.43000)	-0.39033	1.00277	19.82029 (0.66000)
Raw milk-Emmenthal cheese	0.94782 (0.33340359)	-0.22714	1.27432	16.36085 (0.84000)	-0.22714	1.27432	16.36085 (0.84000)
Raw milk-Fresh cheese	1.22849 (0.27124208)	-0.73343	1.27389	33.63615 (0.07000)	-0.32534	1.02475	27.69132 (0.27000)
Raw milk-Manchego cheese	1.29664 (0.25845128)	-1.16843	0.25055	21.59255 (0.66000)	-0.10889	1.26754	16.45669 (0.91000)
Raw milk-Pasteurized milk	1.78807 (0.18520280)	-0.88133	0.12056	38.83688 (0.01000)	-1.03405	1.53312	20.52416 (0.61000)
Raw milk-Cheese in portions	0.55066 (0.46036419)	-0.73461	1.38274	39.61200 (0.00000)	-0.49470	0.76751	34.18943 (0.05000)
Raw milk-Powdered milk	1.43795 (0.23424573)	-0.19895	1.02666	54.52352 (0.00000)	-0.19895	1.02666	54.52352 (0.00000)
Raw milk-Sterilized milk	0.75284 (0.38834734)	-1.33877	0.62637	19.07021 (0.74000)	-1.43692	0.46667	18.59340 (0.75000)
Raw milk-Yogurt	1.39284 (0.24165622)	-0.67333	0.15516	23.24691 (0.46000)	-1.02070	1.27035	17.41636 (0.85000)

Figure 1. Nonlinear Impulse-response functions (monthly data).



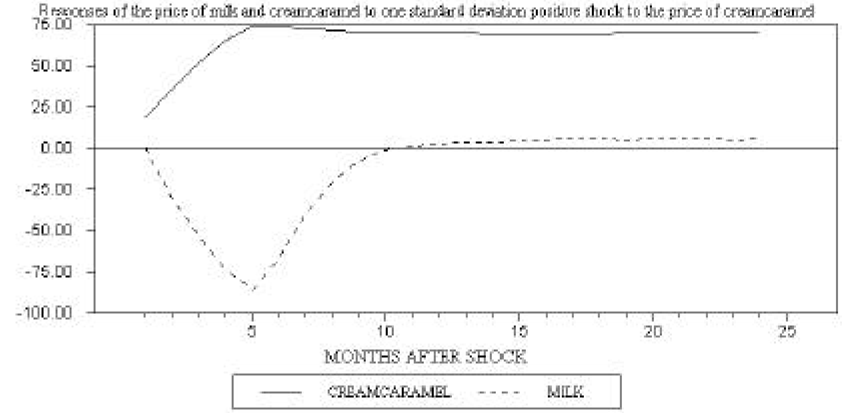
IMPULSE RESPONSE FUNCTION

Responses of the price of milk and continuation milk to one standard deviation positive shock to the price of continuation milk



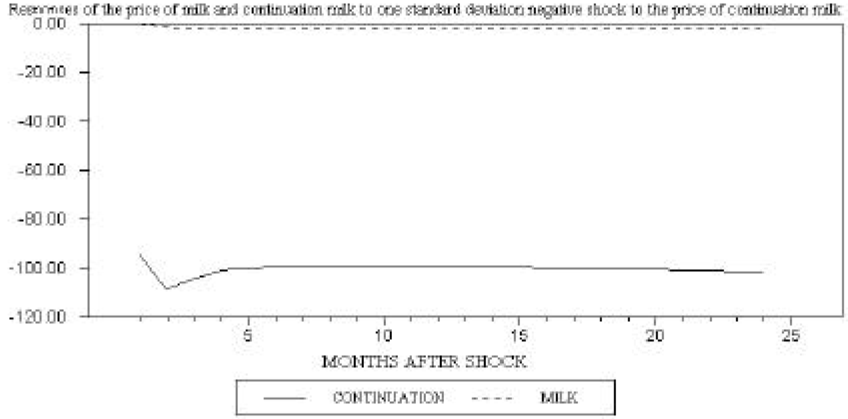
IMPULSE RESPONSE FUNCTION

Responses of the price of milk and creamcaramel to one standard deviation positive shock to the price of creamcaramel



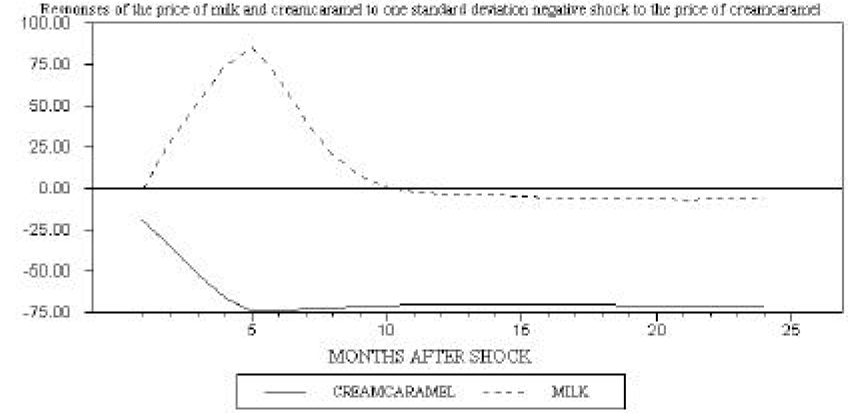
IMPULSE RESPONSE FUNCTION

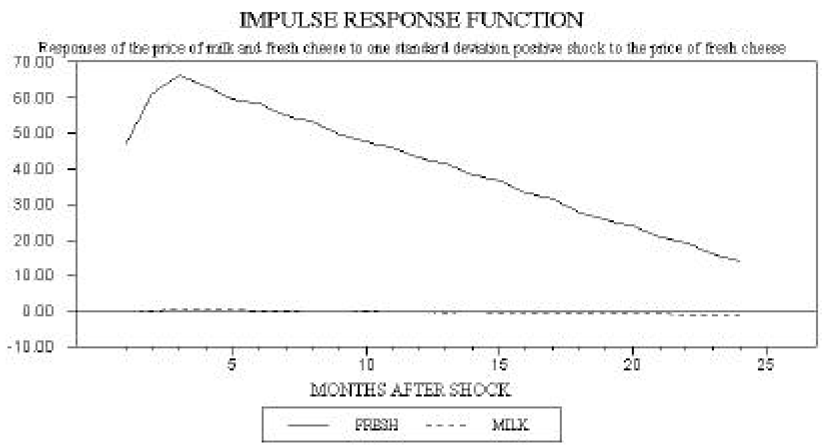
Responses of the price of milk and continuation milk to one standard deviation negative shock to the price of continuation milk

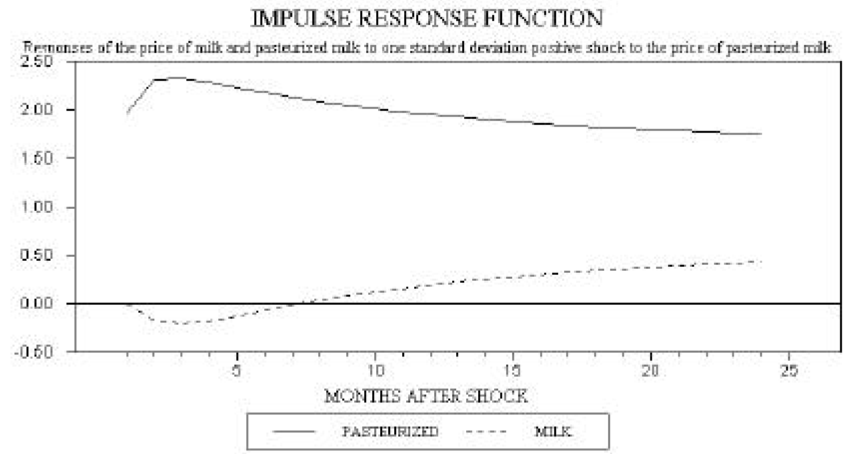


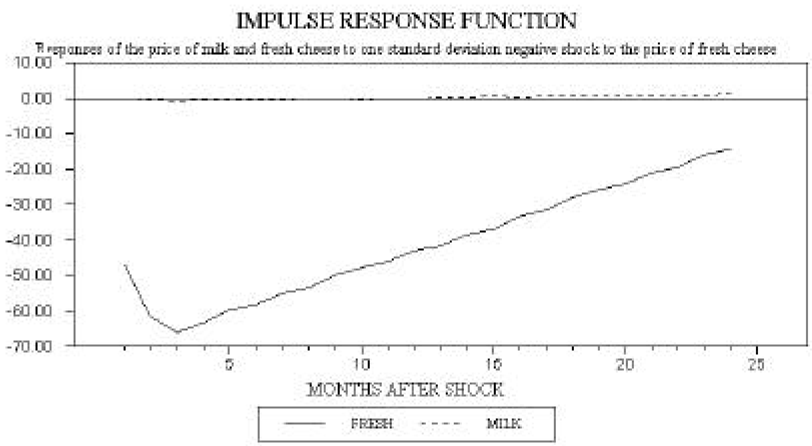
IMPULSE RESPONSE FUNCTION

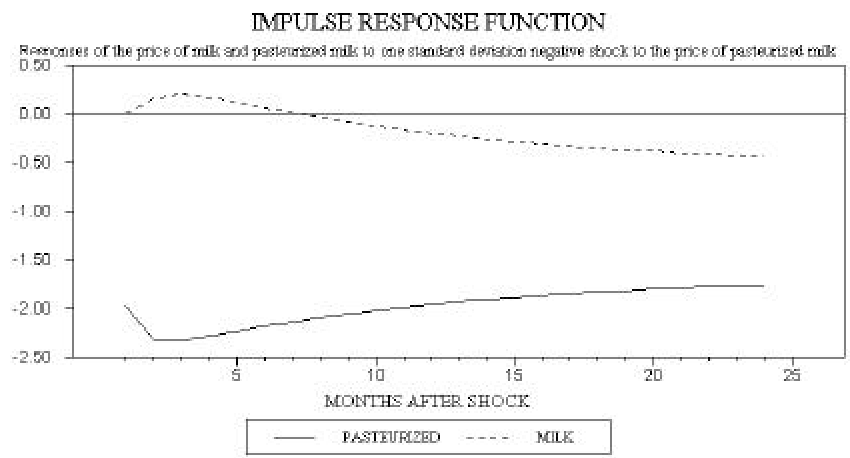
Responses of the price of milk and creamcaramel to one standard deviation negative shock to the price of creamcaramel




IMPULSE RESPONSE FUNCTION
Responses of the price of milk and fresh cheese to one standard deviation positive shock to the price of fresh cheese
70.00
60.00
50.00
40.00
30.00
20.00
10.00
0.00
-10.00
5
10
15
20
25
MONTHS AFTER SHOCK
FRESH
MILK

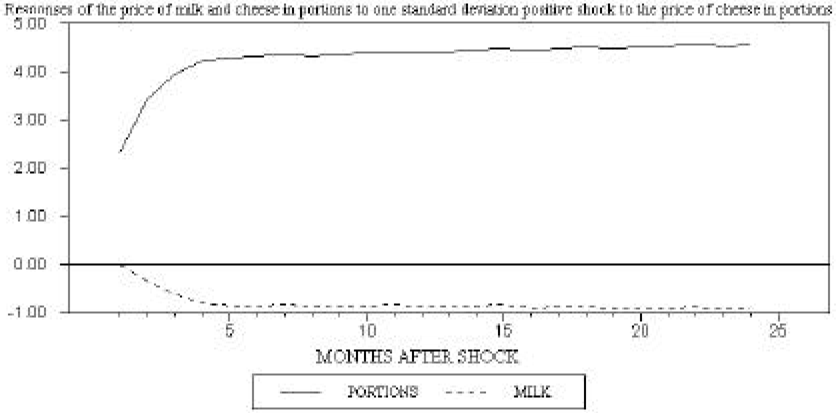

IMPULSE RESPONSE FUNCTION
Responses of the price of milk and pasteurized milk to one standard deviation positive shock to the price of pasteurized milk
2.50
2.00
1.50
1.00
0.50
0.00
-0.50
5
10
15
20
25
MONTHS AFTER SHOCK
PASTEURIZED
MILK


IMPULSE RESPONSE FUNCTION
Responses of the price of milk and fresh cheese to one standard deviation negative shock to the price of fresh cheese
10.00
0.00
-10.00
-20.00
-30.00
-40.00
-50.00
-60.00
-70.00
5
10
15
20
25
MONTHS AFTER SHOCK
FRESH
MILK


IMPULSE RESPONSE FUNCTION
Responses of the price of milk and pasteurized milk to one standard deviation negative shock to the price of pasteurized milk
0.50
0.00
-0.50
-1.00
-1.50
-2.00
-2.50
5
10
15
20
25
MONTHS AFTER SHOCK
PASTEURIZED
MILK

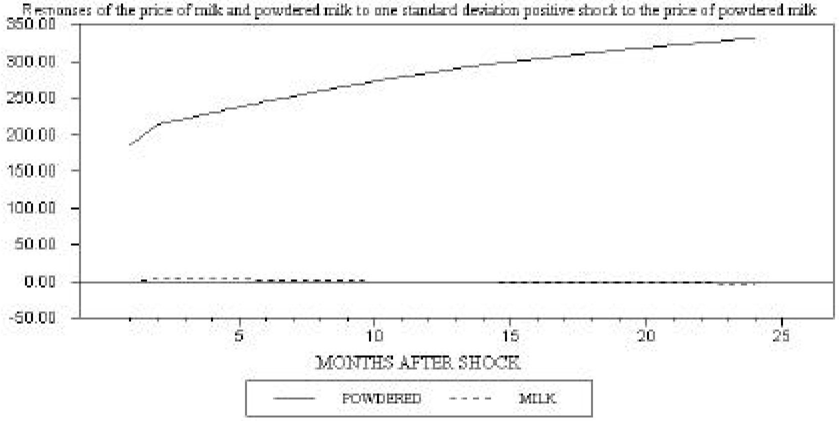
IMPULSE RESPONSE FUNCTION

Responses of the price of milk and cheese in portions to one standard deviation positive shock to the price of cheese in portions



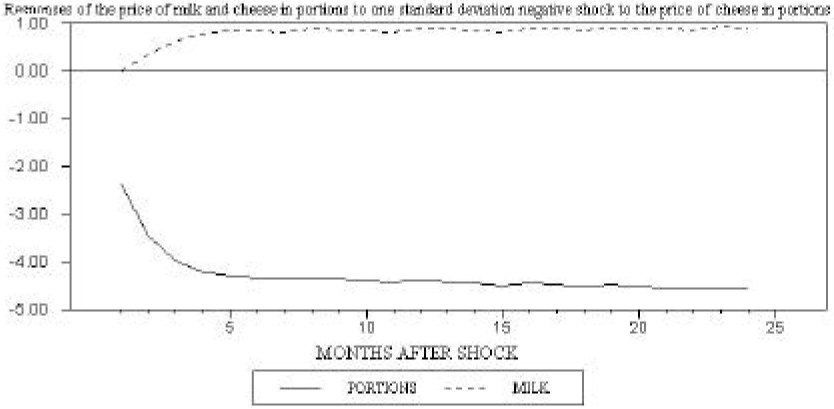
IMPULSE RESPONSE FUNCTION

Responses of the price of milk and powdered milk to one standard deviation positive shock to the price of powdered milk



IMPULSE RESPONSE FUNCTION

Responses of the price of milk and cheese in portions to one standard deviation negative shock to the price of cheese in portions



IMPULSE RESPONSE FUNCTION

Responses of the price of milk and powdered milk to one standard deviation negative shock to the price of powdered milk

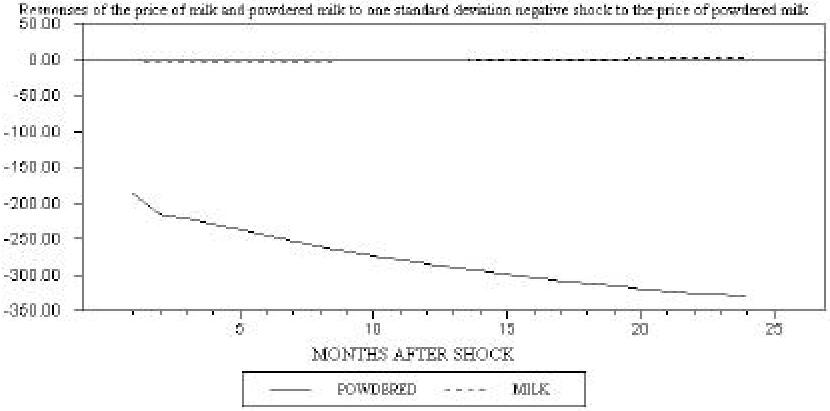
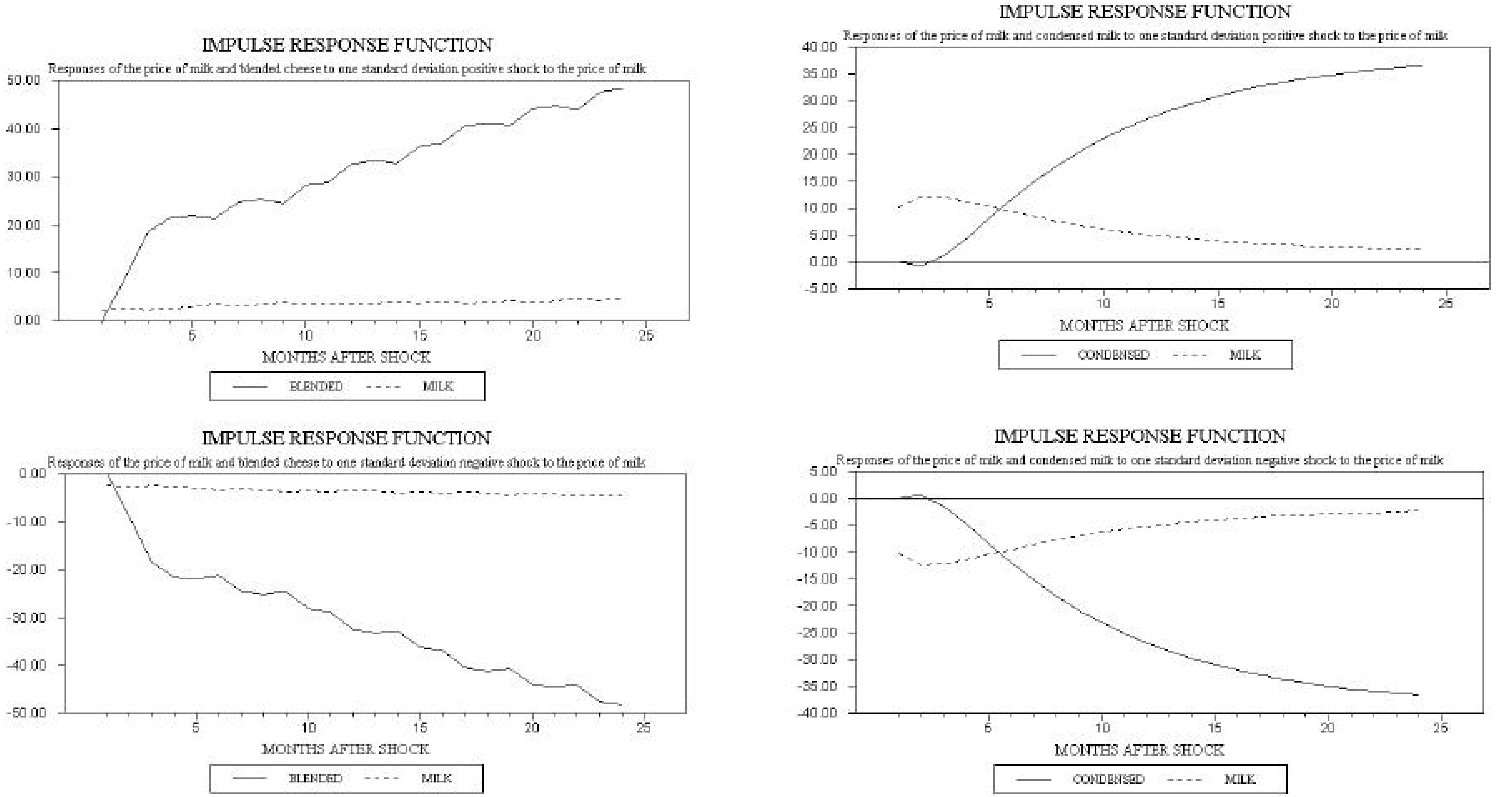
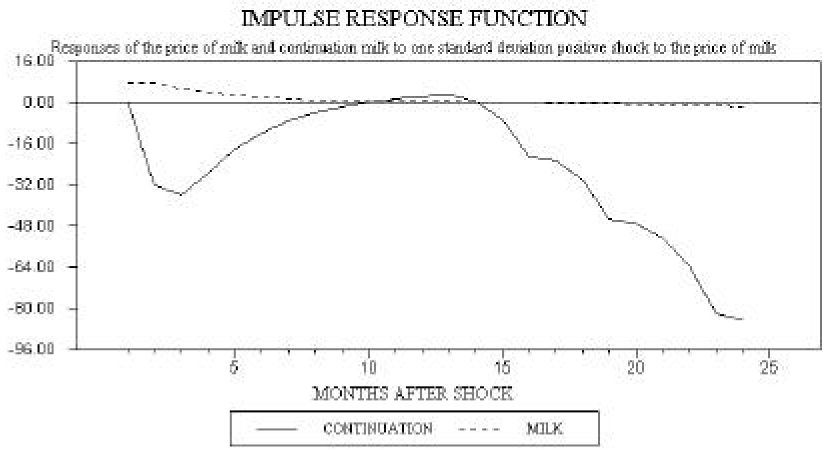
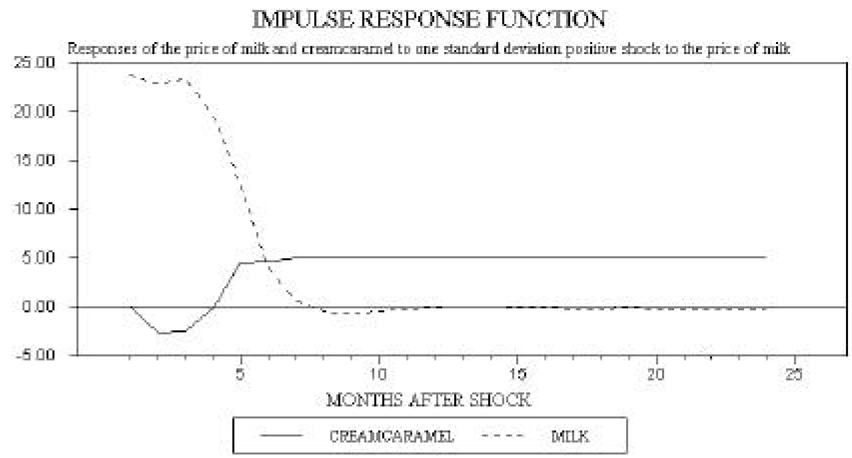
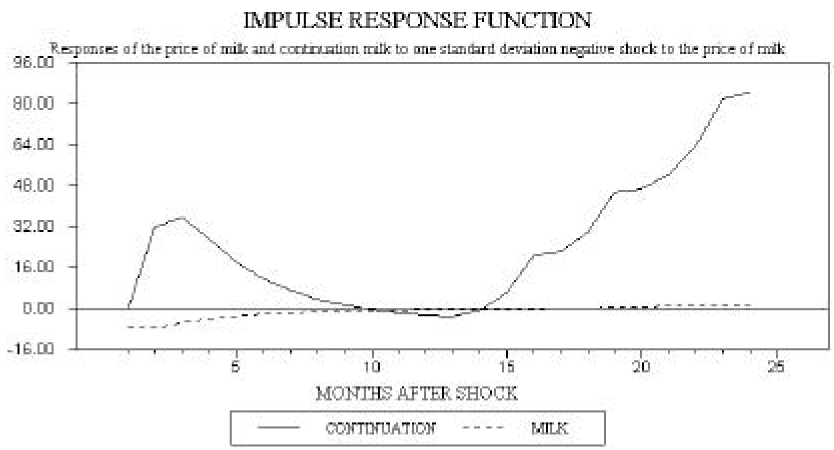


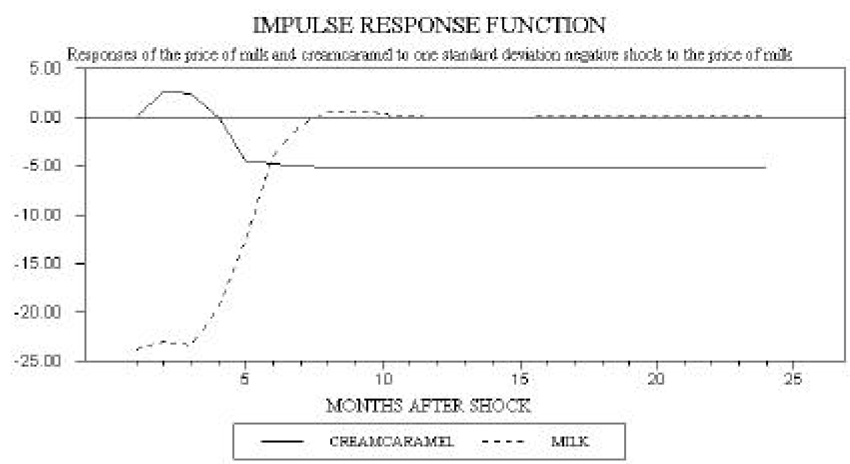
Figure 2. Nonlinear Impulse-response functions (monthly data).

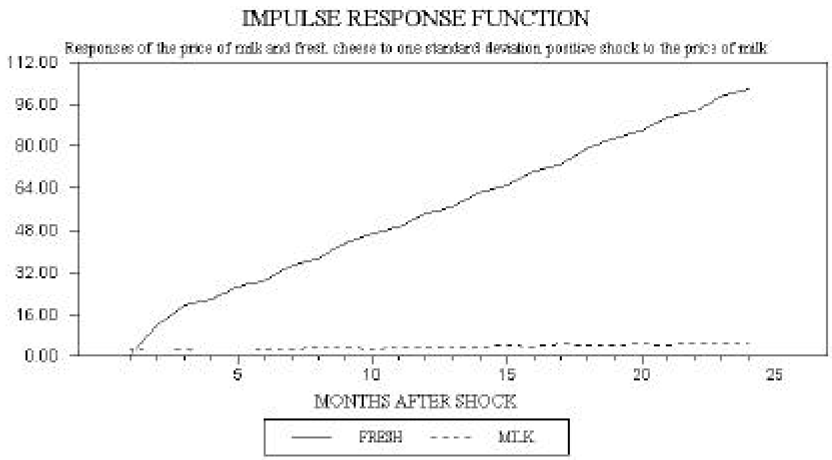


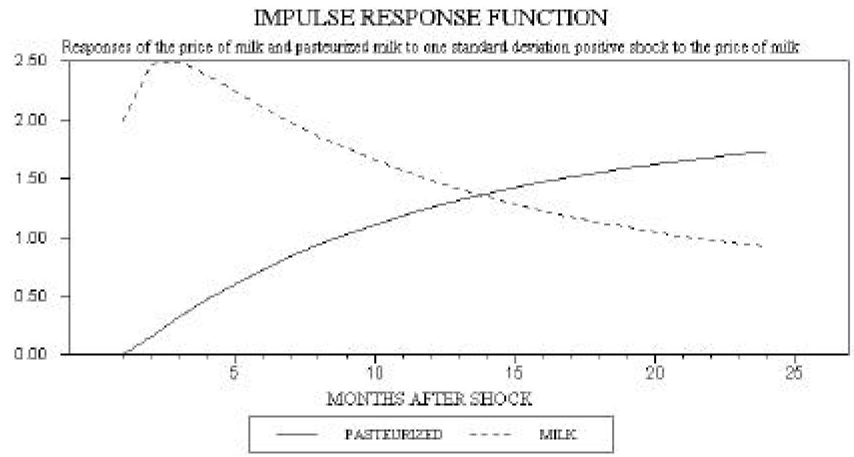

IMPULSE RESPONSE FUNCTION
Responses of the price of milk and continuation milk to one standard deviation positive shock to the price of milk
MONTHS AFTER SHOCK
CONTINUATION
MILK

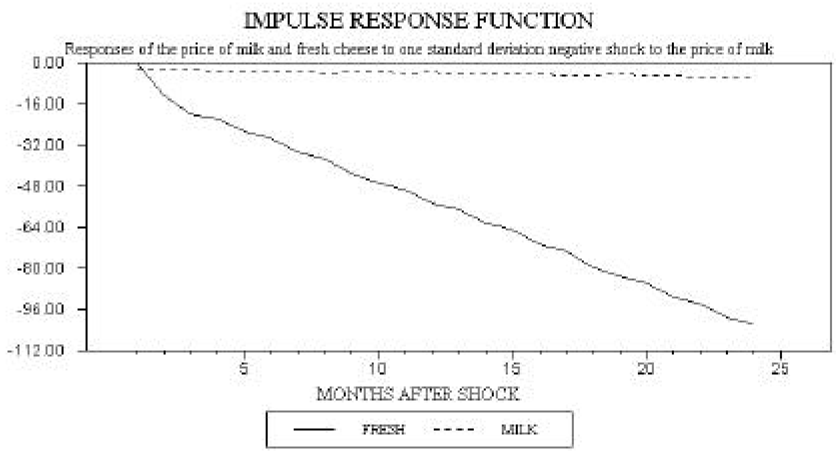

IMPULSE RESPONSE FUNCTION
Responses of the price of milk and creamcaramel to one standard deviation positive shock to the price of milk
MONTHS AFTER SHOCK
CREAMCARAMEL
MILK

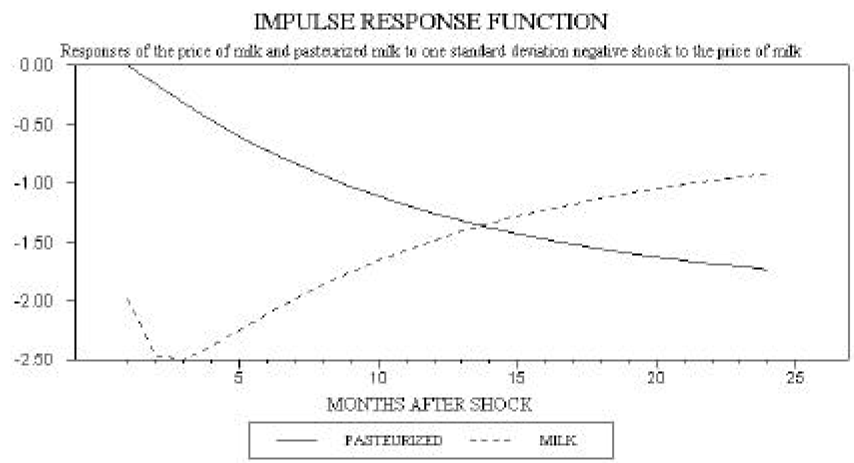

IMPULSE RESPONSE FUNCTION
Responses of the price of milk and continuation milk to one standard deviation negative shock to the price of milk
MONTHS AFTER SHOCK
CONTINUATION
MILK


IMPULSE RESPONSE FUNCTION
Responses of the price of milk and creamcaramel to one standard deviation negative shock to the price of milk
MONTHS AFTER SHOCK
CREAMCARAMEL
MILK

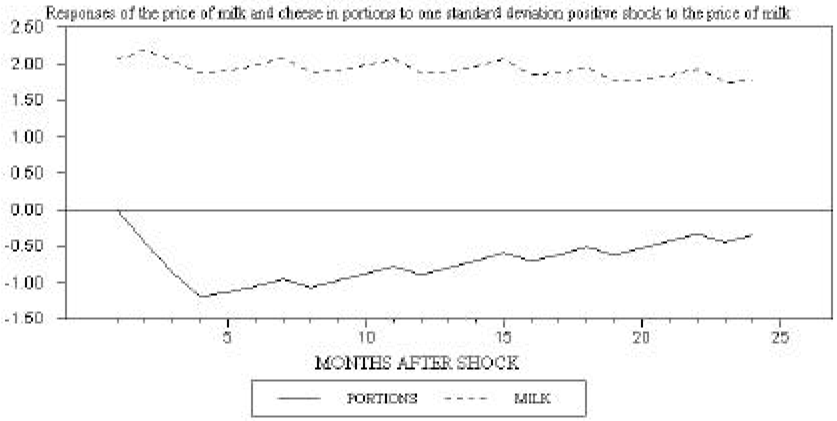

IMPULSE RESPONSE FUNCTION
Responses of the price of milk and fresh cheese to one standard deviation positive shock to the price of milk
112.00
96.00
80.00
64.00
48.00
32.00
16.00
0.00
5
10
15
20
25
MONTHS AFTER SHOCK
FRESH
MILK


IMPULSE RESPONSE FUNCTION
Responses of the price of milk and pasteurized milk to one standard deviation positive shock to the price of milk
2.50
2.00
1.50
1.00
0.50
0.00
5
10
15
20
25
MONTHS AFTER SHOCK
PASTEURIZED
MILK

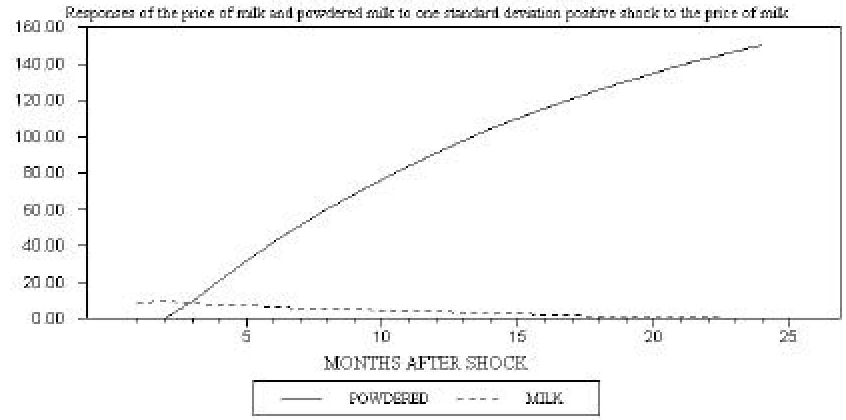

IMPULSE RESPONSE FUNCTION
Responses of the price of milk and fresh cheese to one standard deviation negative shock to the price of milk
0.00
-16.00
-32.00
-48.00
-64.00
-80.00
-96.00
-112.00
5
10
15
20
25
MONTHS AFTER SHOCK
FRESH
MILK


IMPULSE RESPONSE FUNCTION
Responses of the price of milk and pasteurized milk to one standard deviation negative shock to the price of milk
0.00
-0.50
-1.00
-1.50
-2.00
-2.50
5
10
15
20
25
MONTHS AFTER SHOCK
PASTEURIZED
MILK

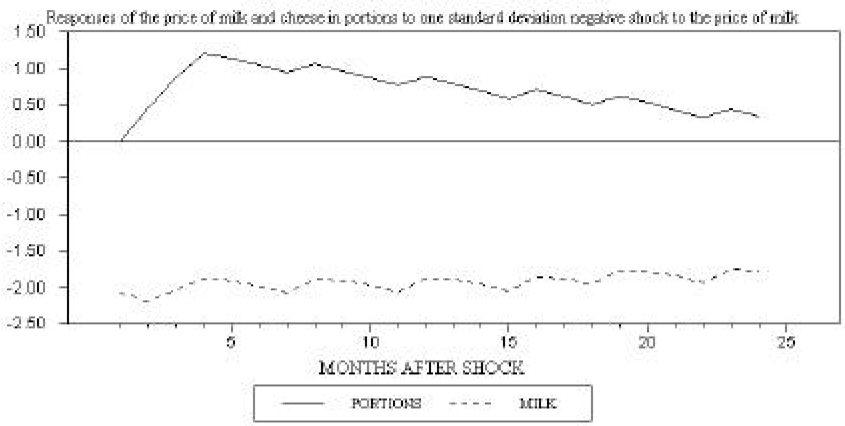
IMPULSE RESPONSE FUNCTION



IMPULSE RESPONSE FUNCTION



IMPULSE RESPONSE FUNCTION



IMPULSE RESPONSE FUNCTION