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Oligopoly and Price Transmission in Turkey's Fluid Milk Market

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ABSTRACT

Farmers and consumers suspect that processing firms abuse their power in the milk marketing chain by engaging in price fixing behavior. The author employs threshold autoregressive and moment threshold autoregressive tests, and contrary to expectations, finds evidence for a downward trend in wholesale milk price without a corresponding decline in farm-gate prices. The downward trend coincides with increased competition in the dairy industry and with the growing market share of the formal sector at the expense of the informal sector. Major dairy processing firms expand their market share and yet continue to enjoy healthy profits thanks to increasing returns due to economies of scale in their processing and distribution operations in a growing market. [JEL Classifications: Q110, Q130].© 2013 Wiley Periodicals, Inc.

1. INTRODUCTION

Two complaints are commonly heard in Turkey regarding milk markets: one among milk farmers, the other among milk processors. Dairy farmers in Turkey complain that farm-gate milk prices in Turkey are relatively low compared to the European Union (EU), whereas the retail price of milk would be among the highest. The Cattle Breeders Central Association complains of collusion among milk processors who seek to avoid bidding against each other in the quarterly auctions at which milk prices are set (Güngör, 2006), but their biggest complaints are low and volatile milk prices. Producing milk suitable for delivery to modern milk processors requires hefty investments in milking machines, cold storage, and high-yielding cow breeds. To recoup such costly investments, farmers need market predictability that extends beyond the immediate quarter. Yet currently there is no national policy in Turkey protecting milk farmers from price fluctuations, whether those fluctuations are a result of ordinary market forces, or engineered by oligopolistic buyers.

On the other hand, milk processors in Turkey complain about the low quality of milk produced in Turkey (i.e., very high bacteria count) and small size of dairy operations, which hampers the ability of milk farmers to modernize their operations. Several sources estimate that only 30% of milk produced in Turkey is processed by modern enterprises (Voorbergen, 2004); the rest of the output is of such low quality that it would not qualify for support under current EU Common Agricultural Policy regulations (Perakende, 2007). The complaints of milk processors, and the small size and dispersed nature of dairy farms in Turkey, are well documented (Food and Agriculture Organisation [FAO], 2007). However, interactions along the dairy marketing chain have been much less studied in Turkey, despite the rich literature on this subject in the United States and Europe (on the U.S. dairy sector, see Capps & Sherwell, 2007; Carman & Sexton, 2005; Chidmi, Lopez, & Cotterill, 2005; Cotterill, 2005; Lass, 2005; Li, 2008). The focus of this article is on whether dairy processing firms abuse their market power to increase profits at the expense of others in the dairy marketing chain.

A likely avenue leading to higher profits for oligopolistic food processors is price manipulation. As a result, one means of investigating the competitiveness of agricultural markets is to study price transmission from the farmers (the primary producers) to the final consumers. Standard economic models assume that positive and negative price changes in input costs are equally transmitted to output prices. However, empirical studies challenge this assumption. Peltzman (2000)–analyzing a wide range of industries—finds that output prices rise concurrently with input price increases (i.e., without lag), but respond with a lag when input

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prices decline. The differing transmission of input price changes to output prices is called *asymmetric price transmission* (APT).

The APT model cannot distinguish between the potential causes of price transmission; instead, the model serves a role in detecting asymmetry in price transmission. According to Meyer and von Cramon-Taubel (2004), the two main theories to explain observed APT are (a) abuse of market power, and (b) adjustment costs (e.g., menu costs in the presence of inflation). Farmers and consumers—those at the beginning and end of the marketing chain, respectively—often suspect abuse of market power in cases of APT. Increases in input prices are passed along more quickly because they squeeze the gross margins of processing firms (the middlemen). Alternatively, asymmetry could be explained by the fact that inflation eats into retail prices; thus, input price declines are not immediately followed by output price declines because the latter result in the reestablishment of previous gross margins that were eroded by inflationary pressures.

Moreover, the mere existence of oligopolies does not necessarily imply anticompetitive behavior, and sometimes the benefits of oligopolies outweigh their potential costs. For example, oligopolies may enjoy "super profits" (in comparison to the perfect competition case) yet deliver lower prices to consumers because they enjoy economies of scale. Alternatively, if barriers to entry are sufficiently low, the threat of entry can force existing monopolies or oligopolies to behave as if they were operating in a competitive market. Hence, to determine the welfare effects of oligopolistic market structure, it is not enough to measure the market share of larger firms. Indeed, McCorriston, Morgan, and Rayner (2001) provide a theoretical framework wherein increasing returns to scale in oligopolistic markets can lead to even greater price transmission than a perfect competition case:

> Specifically, whereas market power will reduce the level of price transmission (relative to perfectly competitive case), if the industry is characterized by increasing returns to scale, the level of price transmission will increase. Under reasonable conditions, the degree of price transmission may be greater than in the constant returns, perfectly competitive case. (p. 146)

The framework suggested by McCorriston et al. (2001) offers a fruitful approach to investigate the fluid milk market in Turkey, which both demonstrates falling prices for processed milk and concentration of production among a small number of firms (five: SEK, Pınar, Sütaş, Danone, and Ülker), i.e., a seemingly oligopolistic market wherein gross margins between input prices (farm-gate milk price), and output prices (ultra-high temperature [UHT] milk price) that have narrowed instead of widening. The most approximate explanation of the functioning of the fluid milk production and processing chain is that milk processors enjoy oligopsony powers, and hence can extract price concessions from the farmers. However, processing firms are passing along price concessions and more, to the retailers[1] because they also enjoy increasing returns to scale. Increasing returns to scale allows dairy firms to preserve their net profit rates even though the gross margin between the farm-gate milk price (the chief input) and UHT milk is narrowing.

This paper utilizes a two-pronged approach to test empirically the McCorriston et al., (2001) framework in Turkey: first, I include an hourly labor productivity index in the dairy sector (hereafter the labor productivity index) on the right-hand side of cointegrating equations (both for UHT and farm-gate milk prices) to account for long-term changes in returns to scale. Second, two major firms (Danone and Ülker) entered the fluid milk market in Turkey in late 1997, so I construct a dummy variable to include in the cointegrating regression to account

[1] McCorriston (2002) points out that not only food processing but also food retailing in Europe is oligopolistic in nature. However, Celen, Erdoğan, and Taymaz (2005) assert that the supermarkets behave competitively in Turkey.

for their entry into the market.[2] Structural break tests for the UHT milk price and for the labor productivity index suggest October 1998 for the break date in the series. Furthermore, available evidence from the Ministry of Agriculture and Rural Affairs (MARA, 2004) shows that the capacity utilization ratio in modern dairy industry increased in Turkey in most of the study period (1994–2001), but also that this remains low. I believe that the availability of excess capacity makes it possible that increasing returns to scale is possible even in the short run. I also find direct evidence for increasing returns to scale: There is also a structural break in unit root tests for the labor productivity index coinciding with the entry of Danone and Ülker into the dairy market in 1997.

Finally, I employ threshold autoregressive (TAR) and moment threshold autoregressive tests (M-TAR) to look for empirical evidence of abuse of market power by milk processing firms. I do not find much evidence for asymmetry, and the little evidence I find for asymmetry supports the contrary conclusion that milk processing firms in Turkey in the study period more quickly transmit input price decreases to retailers than input price increases.

The rest of the article is organized as follows: In the next section unit root tests are described, which are required for any study of time series data. Then the TAR and M-TAR tests are introduced, which are used to test for cointegration and asymmetric price transmission. The results are presented in Section 3 and the conclusions in Section 4.

2. MATERIALS AND METHODS

2.1. Data

Unfortunately, the retail milk price data released by State Institute of Statistics (SIS) includes both open-sourced milk and packaged milk, and hence does not correspond one-to-one to the wholesale product (UHT packed milk) of the modern dairy industry. Hence, this paper concentrates on the relationship between farm-gate (input) and UHT wholesale prices (output). Monthly price data are available from January 1994 on the SIS website http://www.tuik.gov.tr/Start.do (all the price data is indexed using monthly wholesale price index to omit the effects of inflation). I use available data up until to the end of 2006.[3] I want to control for returns to scale while exploring the relationship between farm-gate and wholesale milk prices, but unfortunately, capacity utilization rate data are only available annually. Hence, I use the labor productivity index as a proxy for returns to scale. This index is available only as quarterly data for the period, so I convert the quarterly index data into monthly data by interpolating.

Figure 1 shows that gross margins between indexed farm-gate milk price and indexed wholesale UHT prices are narrowing; Figure 2 shows that the labor productivity index trends upward after 1997. Figure 3 presents time graph indices of three indexed variables, which illustrate the fluctuating movement of farm-gate prices, the long-term decline in UHT prices, and the gradual increase in the productivity index. On its own, Figure 1 suggests that the two series may not be cointegrated in the long-run since the relationship between the two series is not constant. However, taken together, the increase in the productivity index can account as the link between fluctuating farm-gate prices and the long-term decline in UHT wholesale prices.

[2]Structural break tests (more on those tests in Section 2.2) independently confirm that right around the same time UHT milk prices started their long-term decline.

[3]The analysis ends at the end of 2006 because the labor productivity index for the dairy sector is available only until the end of 2006. Recently, the SIS has released data for indices for labor hours and total dairy production for the post-2006 period, but the base year has been changed from 1997 to 2005, so they are not comparable. Furthermore, from early 2007 until mid-2009, the increase in feed prices outpaced milk and beef prices by a wide margin, leading to widespread culling of dairy herds and substantial changes in milk production in Turkey. Hence, it is probably healthier to limit the time as it is here to focus on the behavior of dairy processing firms.

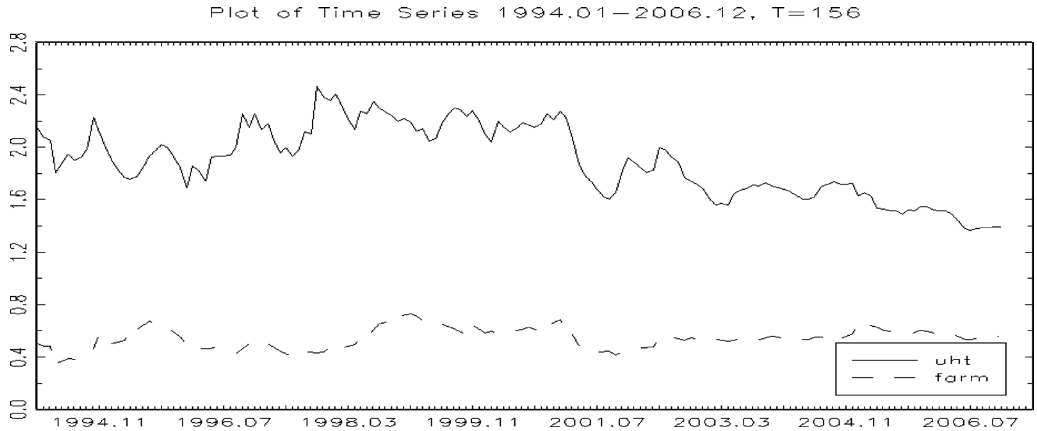

Figure 1 Inflation Adjusted Farm-Gate and Ultra-High Temperature (Wholesale) Milk Prices, YTL. YTL = 1,000,000 Turkish Lira, SIS website http://www.tuik.gov.tr/VeriTabanlari.do?ust_id=6&vt_id=16.

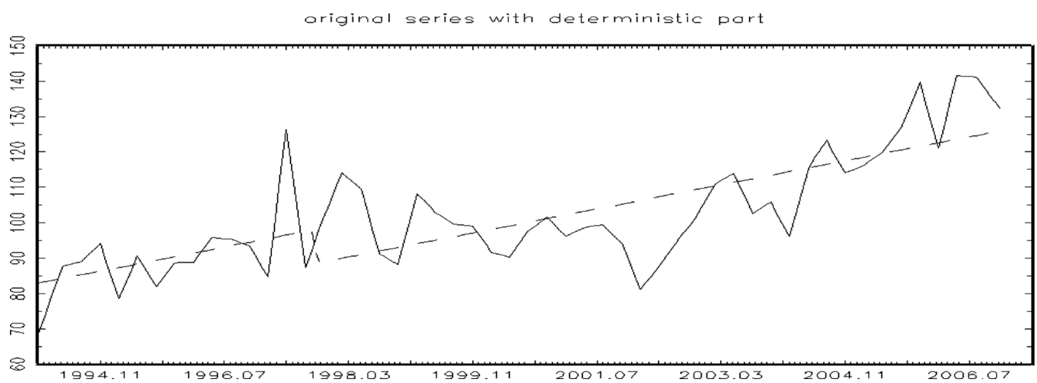

Figure 2 Productivity Index With Shift Dummy, Break (1997.M10), 22 Lags.

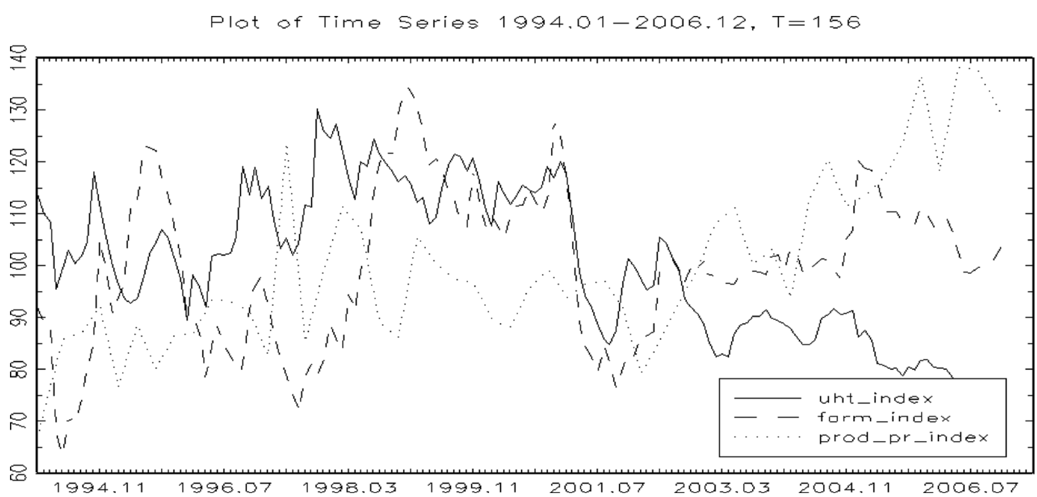

Figure 3 Inflation Adjusted Farm-Gate and Ultra-High Temperature Milk Prices, and Productivity Index.

2.2. Unit Root Tests

After visual analysis of time-series data, I check for the existence for unit roots in time-series variables.[4] Our testing strategy for unit roots is to start with an augmented Dickey-Fuller (ADF) model with one lag including the trend variable, if the trend is visible in the data. If the final prediction error (FPE) score indicates a lag length other than the default one lag, I conduct the ADF test with the suggested lag length. Finally, because the data is monthly, I add monthly dummy variables to see if the results changed significantly. In the second step, I choose the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) test (Kwiatkowski, Phillips, Schmidt, & Shin, 1992) as an alternative to the augmented Dickey–Fuller test (ADF) because it tests for the opposite null hypothesis. In the ADF test, the null hypothesis is the existence of a unit root. In KPSS, the null hypothesis is stationarity. While testing for both level stationarity and trend stationarity, I conduct KPSS tests with the same lag length as the ADF tests. In the last step, I test for the presence of a unit root with a structural break, as suggested by Saikkonen and Lütkepohl (2002). They suggest first estimating Equation (1) and subtracting that from the original series. Then the ADF test is performed on the adjusted series.

$$y_t = \alpha_0 + \alpha_1 t + \gamma D_s + e_t \tag{1}$$

where $Ds = 0$ if $t < T$

$Ds = 1$ if $t \geq T$ and T is the shift date. y_t is the time series, t stands for time and D_s is the dummy variable defined as above.

2.3. Threshold Autoregressive and Moment Threshold Autoregressive Tests

Once it is established that each series has unit root characteristics, the next step is to estimate cointegrating equations. However, the conventional test for cointegration— the Johansen trace test—is known to function poorly when applied to data with asymmetric transmission rates. Two alternative tests have been developed to improve the cointegration test when asymmetric transmission is suspected (Enders, 2004): threshold autoregressive (TAR) and moment threshold autoregressive (M-TAR). To perform these alternative tests, we first need to estimate the long-term relationship between UHT real prices (*uht*) and farm-gate real prices (*farm*) to obtain the residuals. Following the findings in unit root tests, I include the productivity index ($prod_t$), and the trend term (t) on the right-hand side in Equations (2a) and (2b). Furthermore, I add a corresponding structural break dummy variable suggested by the findings from the unit root tests (a shift dummy variable from October 1997 onwards for UHT milk (DV9710), and a shift dummy variable for the time period from December 2000 onwards for farm-gate milk real price (DV0012)).

$$uht_t = \alpha_1 + \beta_{u1} t + \beta_{u2} farm_t + \beta_{u3} prod_t + \beta_{u4} DV9710 + \mu_t \tag{2a}$$

$$farm_t = \alpha_2 + \beta_{f1} t + \beta_{f2} uht_t + \beta_{f3} prod_t + \beta_{f4} DV0012 + \theta_t \tag{2b}$$

$$\text{where } \mu_t = \rho\mu_{t-1} + \varepsilon_t \text{ and } \theta_t = \eta\theta_{t-1} + \tau_t \tag{3}$$

In the TAR model, the coefficients of lagged error correction term, μ_t and θ_t, are allowed to take different values across a threshold (Enders & Siklos, 2001):

$$\begin{aligned} \Delta\mu_t &= I_t \rho_1 \mu_{t-1} + (1 - I_t)\rho_2 \mu_{t-1} + e_t \\ I_t &= \begin{cases} 1 \ if \ \mu_{t-1} \geq c \\ 0 \ if \ \mu_{t-1} < c \end{cases} \end{aligned} \tag{4}$$

[4] We perform all unit root tests with J-Multi software (http://www.jmulti.de/)

If c is equal to zero, μ_{t-1} is a simple negative or positive deviation from equilibrium. We expect ρ_1 and ρ_2 to be negative so that deviations adjust toward the long-run equilibrium. If the deviation of the UHT milk price from long-run equilibrium is positive (more generally greater than c) in the previous period, then $\rho_1\mu_{t-1}$ will be eliminated in the current period, and vice versa. The values of ρ_1 and ρ_2 indicate the relative speed of adjustment. If $\rho_1 > \rho_2$, faster convergence is observed when the prices are above the equilibrium.[5]

The second alternative framework accommodating asymmetry is the M-TAR model:

$$\begin{gathered}\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + e_t \\ I_t = \begin{cases} 1 \; if \; \Delta\mu_{t-1} \geq c \\ 0 \; if \; \Delta\mu_{t-1} < c \end{cases}\end{gathered} \tag{5}$$

In the case of the M-TAR model, economic agents adjust their behaviors according to the trend, or "momentum," of deviations instead of adjusting their behavior according to the deviations themselves. In other words, ρ_1 and ρ_2 describe adjustments in response to momentums in different directions. If $\rho_1 \neq \rho_2$, the adjustments are not symmetric and shows more "momentum" in one direction than in the other.

Following Enders and Siklos (2001), I perform a grid search to determine the value of the threshold. After sorting all of the estimated $\mu_t(\Delta\mu_t)$ from Equation (2a) in ascending order, we consider values between the 15th percentile and 85th percentile as possible threshold values. These values are used to estimate Equations (4) and (5). The value that yields the lowest residual sum of squares is deemed to be the appropriate threshold value.

To ensure cointegration, ρ_1 and ρ_2 should be negative so that the long-term relationship between the variables do not deviate or shrink. The negative coefficients ensure that the short-term deviations are corrected towards the long-term equilibrium. Enders and Siklos (2001) obtained critical values by recording the *t* statistics for the two null hypotheses $\rho_1 = 0$ and $\rho_2 = 0$ and the *F* statistic for the joint hypothesis $\rho_1 = \rho_2 = 0$. In the *t* tests, the larger of the two *t* statistics is called *t-Max*, and the smaller is called *t-Min*. If series are cointegrated ρ_1, ρ_2, and the corresponding *t* statistics, should be negative (*t-Min* < *t-Max* < 0). In the *F* test, the *F* statistic for the joint hypothesis of $\rho_1 = \rho_2 = 0$ is called Φ, to distinguish it from the usual *F* distribution. According to Enders and Siklos (2001), the Φ statistic has substantially more power than *t-Max* and *t-Min* statistics. Hence, for brevity, I only discuss the more powerful Φ statistic in the results section. Finally, Enders and Siklos (2001) report that the M-TAR test is more powerful than the TAR test.[6]

3. RESULTS

3.1. Unit Root Tests

For the sake of brevity, I present only unit root test results with lag lengths minimizing the FPE score. For the UHT milk real price and labor productivity index, I present unit root tests, taking into account the prior information as there is a visible trend in the time series (Elder & Kennedy, 2001). Table 1 shows that both the ADF and KPSS tests gave the same results for the UHT milk real price: The monthly prices are nonstationary. However, the conclusions differ when I conduct the unit root test with a structural break test. The structural break test suggests a break date of October 1997. The most significant developments prior to or during 1997 were

[5] If $y_t > \hat{y}_t$ then μ_t will be positive. In other words, because $\hat{y}_t$ is the long-run equilibrium, positive μ_t indicates that actual price is above the long-run price.

[6] In other words, the TAR test correctly rejects the null hypothesis of no cointegration less often than the Engle-Granger methodology in Monte Carlo experiments. Enders and Siklos (2001) suggest that in the case of the TAR model, the gain from estimating the correctly specified model (asymmetric) is outweighed by the estimation of an additional coefficient–threshold (p. 171).

TABLE 1. Unit Root Tests for Ultra-High Temperature (UHT) and Farm-Gate Milk Prices and Labor Productivity Index

Variable	Test	Structural break date	Trend variable	Lags	Test score	Conclusion
UHT milk real price	DF		Yes	13 Lags	−1.5562	FTR Ho of unit root
	KPSS		Yes	13 Lags	0.2455	Reject Ho of stationarity
	Structural break	1997 M10	Yes	3 Lags	−4.0588	**Reject Ho of unit root**
farm-gate milk real price	ADF		Yes	1 Lag	−2.7427	FTR Ho of unit root
	KPSS		Yes	1 Lag	0.271	Reject Ho of stationarity
	Structural break	2000 M12	Yes	1 Lag	−2.6061	FTR Ho of unit root
productivity index	ADF		Yes	22 Lags	−1.1158	FTR Ho of unit root
	KPSS		Yes	22 Lags	0.1407	Reject Ho of stationarity
	Structural break	1997 M10	Yes	22 Lags	−1.6154	FTR Ho of unit root

Note. KPSS = Kwiatkowski–Phillips–Schmidt–Shin; ADF = augmented Dickey–Fuller test; FTR = fail to reject.

TABLE 2. Long-Term Relationship for Inflation-Adjusted Ultra-High Temperature (UHT) Milk Price

Dependent	UHT milk	Farm-gate milk
UHT		0.11550***
Farm-gate	0.40710**	
Time trend	(0.00795)***	0.00130***
Productivity index	0.00005	(0.00017)
Str. break DV (1997 M10)	0.46588***	
Str. break DV (2000 M12)		(0.03370)
Constant	5.19804***	(0.27723)
Adj. R^2	0.7023	0.1924

*** 1%. **5%.

the privatization of SEK (the publicly owned milk company)[7] and entrance of new firms into the dairy sector. The privatization of the SEK was for the most part completed during August–September 1995. French multinational Danone bought a local company, Tikvesli, and entered the Turkish market in 1997. Again in 1997, the largest domestic food company, Ülker, entered the consumer dairy market (Voorbergen, 2004).[8] In other words, the structural break identified in empirical analysis coincides with a significant change in the structure of the industry.

Table 1 also shows that the ADF and KPSS for farm-gate milk real price and labor productivity index, and supports the unit root null hypotheses in every alternative test. The structural break date is December 2000 for farm-gate milk real price. In the case of the labor productivity index, the structural break date is October 1997. As discussed above, major new firms entered the dairy sector during 1997. Figure 2 shows the positive deterministic time trend imposed on actual productivity index data.

3.2. Long-Term Relationship

Table 2 presents the estimates for long-term relationships as depicted in Equations (2a) and (2b). The equation for the UHT price explains more than 70% of the variation in real price of UHT milk in the study period. The coefficient estimate for the farm-gate price is positive and significant at 0.4071, and the corresponding long-term price elasticity evaluated at mean

[7] Modern enterprises, private or owned by the state, have never processed the majority of raw milk in Turkey. However, SEK, due to its presence in every region, was widely believed to set the reference price both for farmers (when buying milk from) and for processors (when selling their products).

[8] "Ülker is a diversified Turkish food company with sales of around USD 2.5 billion [2004] that expanded into the dairy business relatively recently. The company already manufactured powdered milk for its own cookie business but moved into the end consumer business with the acquisition of Ak Foods in 1997." (Voorbergen, 2004, p. 9)

TABLE 3. Results of Threshold Autoregressive (TAR) and Moment Threshold Autoregressive Tests (M-TAR) for Inflation Indexed Ultra-High Temperature Milk Price

	Threshold	ρ_1	ρ_1	Φ^a	$\rho_1 = \rho_2{}^b$	p Value
TAR						
c = 0		(0.217)	(0.133)	8.06	0.79	0.38
c ≠ 0	0.131	(0.253)	(0.110)	8.95	2.4	0.12
M-TAR						
c = 0		(0.362)	0.044	19.44	21.49	0.00
c ≠ 0	0.058	(0.445)	0.000	21.89	25.93	0.00

[a] *F* statistics for the joint hypothesis $\rho_1 = \rho_2 = 0$. When c = 0: TAR: 1%: 8.24, 5%: 5.98; 10%: 5.01; M-TAR: 1%: 8.78, 5%: 6.51, 10%: 5.45. When $c \neq 0$: TAR: 1%: 9.27, 5%: 6.95; 10%: 5.95; M-TAR: 1%: 9.14, 5%: 6.78, 10%: 5.73. [b] *F* statistics for the joint hypothesis $\rho_1 = \rho_2$ to test for asymmetric price transmission. The test statistics are taken from Enders and Siklos (2001).

values is 0.116 given average values for UHT and farm-gate real milk prices. In addition to the farm-gate price, both trend and dummy variables are statistically significant (but not the labor productivity index). Both the trend variable and dummy variable significantly improve the fit of the model, and both are economically significant. The effect of dummy variable is equivalent to 25% of the mean value of UHT price (0.466/1.89 = 25% percent) during the period in question; the trend coefficient (−0.00795) is roughly equivalent to 0.1 TL (or 0.1/1.89 = 5%) annual decline in real prices during the study period.

The equation for the farm-gate price explains significantly less, only 20% of the variation in farm-gate real prices during the study period. The UHT milk price and trend variable are the only two statistically significant variables. The coefficient estimate for the UHT price is positive and significant at 0.1155, and the corresponding long-term price elasticity evaluated at mean values is 0.404 at average values for UHT and farm-gate real milk prices. In other words, a 10 percent increase in UHT milk price will result in a slightly more than 4% increase in farm-gate milk price. The trend variable is positive, which is not the case for UHT milk, and at 0.0013, suggests a 0.016 TL (or 0.016/0.54 = 3%) annual increase in farm-gate real prices during the study period.

3.3. TAR and M-TAR Tests for UHT Milk Price

I can confidently claim unit roots for all variables in every case, except when I introduce a structural break into the unit root analysis of the inflation adjusted UHT milk price. Despite this caveat, I continue with the cointegration analysis assuming that all variables are unit-root processes. Table 3 shows the test results for both TAR and M-TAR models for Equation (2a). For the TAR model, all coefficient estimates are negative, as expected, for both zero and nonzero thresholds. I reject the null hypothesis of no cointegration at a 5% significance level with the Φ test. The next step is to test for asymmetry. The usual *F* test is sufficient in this case (the last two columns of Table 2). For the TAR model, we fail to reject the symmetry hypothesis in both cases (zero and nonzero thresholds). Test results show that in the case of the TAR model, the threshold is more than zero, indicating that milk processors adjust prices when the actual wholesale prices are above the long-term equilibrium price. Moreover, the absolute value of the coefficient estimate of ρ_1 is larger than that of ρ_2, suggesting faster convergence in response to positive deviations from equilibrium.

In the M-TAR test, coefficient estimates for ρ_1 are negative, but coefficient estimates for ρ_2 are positive (and not statistically significant) both in the zero and nonzero threshold cases. The sample Φ statistics are 19.49 for zero threshold and 21.89 for nonzero threshold. The Φ statistics are greater than the 1% significance value of 8.85, so the null hypothesis of no cointegration can be rejected. Given these strong results for cointegration, I test for asymmetry. The *F*-test statistics lead to a strong rejection of the null hypothesis of symmetry. In the case of the M-TAR model, the threshold is more than zero (similar to TAR model), indicating that milk processors adjust prices when the deviations from long-term equilibrium are above the long-term for "momentum." Moreover, the absolute value of the coefficient estimate of ρ_1 is larger than

TABLE 4. Results of Threshold Autoregressive (TAR) and Moment Threshold Autoregressive Tests (M-TAR) for Inflation Indexed Farm-Gate Milk Price

	Threshold	ρ_1	ρ_2	Φ^a	$\rho_1 = \rho_2{}^b$	p Value
TAR						
$c = 0^a$		(0.076)	(0.091)	3.35		
$c \neq 0^b$	0.075	(0.063)	(0.102)	3.51		
M-TAR						
$c = 0$		(0.025)	(0.120)	4.40		
$c \neq 0$	0.015	0.023	(0.127)	5.70	4.56	0.034

[a],[b]Same as Table 2.

[a]Tests reveal that residuals for TAR model are not white noise. After augmenting to six lags, we obtained white noise residuals at that point the estimated coefficients became negative, we conclude for cointegration at 5% and symmetry. [b]Tests reveal that residuals for TAR model are not white noise. After augmenting to six lags, we obtained white noise residuals and at that point, the estimated coefficients became negative, we conclude for cointegration at 5% and symmetry.

that of ρ_2, suggesting faster convergence in response to positive deviations from equilibrium (again similar to TAR model). Therefore, the farm-to-wholesale price transmission in Turkey is asymmetric according to both TAR and M-TAR models, and adjustments are stronger when the previous period deviation is positive. That is, when actual wholesale prices are higher than the long-run equilibrium prices, a more rapid adjustment back towards the equilibrium price occurs. In other words, dairy firms tend to be quicker in lowering prices.

3.4. TAR and M-TAR Tests for Farm-Gate Milk Price

I repeat the TAR and M-TAR cointegration analysis with the inflation adjusted farm-gate milk prices as the dependent variable. The results, shown in Table 4, are significantly different. The absolute value of the coefficient estimates for ρ_2 is larger than that of ρ_1, suggesting faster convergence in response to negative deviations from equilibrium. In the case of M-TAR, I fail to reject the no cointegration null hypothesis for the zero threshold. When the threshold is nonzero, the Φ test score is slightly less than the 10 % critical value of 5.73. If I assume the existence of cointegration, then there is evidence for asymmetry at the 5 % confidence level (last two columns). In this case, the threshold is more than zero, indicating that farmers make price adjustments when the deviations from long-term equilibrium are above the long-term momentum. Moreover, the absolute value of the coefficient estimate of ρ_2 is larger than that of ρ_1, suggesting faster convergence in response to negative deviations from equilibrium. Therefore, wholesale-to-farm price transmission in Turkey is asymmetric, and adjustments are stronger when the previous period deviation is negative. That is, when actual farm-gate prices are lower than the long-term equilibrium prices, a more rapid adjustment back toward the equilibrium price occurs. Nevertheless, evidence is weaker when farm-gate price is the dependent variable.

4. CONCLUSION

Time series variables are beset by nonstationarity. I test for the presence of a unit root in inflation-adjusted farm-gate milk prices, UHT milk prices, and the labor productivity index. I find evidence for a structural break during October 1997 for the UHT milk price and labor productivity index, and evidence of another during December 2000 for the farm-gate milk price. Even after accounting for these structural breaks, I find evidence for unit root in most specifications and concluded that these variables are nonstationary. However, these series are stationary in their first differences and hence can be studied in a cointegrated framework where I test for the existence of a long-term relationship between the farm-gate and UHT milk prices. Nevertheless, I suspect asymmetric price transmission and the conventional Johansen trace test is known to perform poorly in the presence of asymmetry. Hence, I apply TAR and M-TAR procedures to test for cointegration in the case of asymmetric price transmission. When the dependent variable is the inflation-adjusted UHT milk price, I find strong evidence for cointegration both with TAR and M-TAR tests.

Interestingly, the asymmetry when the dependent variable is the UHT milk price is the opposite of what most layman predict. The estimated threshold is greater than zero, suggesting that UHT milk producers adjust their prices more quickly when the difference is above the long-run equilibrium (i.e., when gross profit margin is extended). In the M-TAR procedure, I test for whether agents adjust their behaviors according to trends in the deviations instead of adjusting their behavior according to deviations. I find that the absolute value of the ρ_1 (coefficient of deviations that are above the threshold) is larger than ρ_2, meaning that speed of adjustment is greater when the deviations are above the long-run relationship. The combined evidence from the TAR and M-TAR tests for the UHT milk price and the auxiliary evidence of economies of scale support the framework proposed by McCorriston et al. (2001). For UHT milk in Turkey, there is both scope for increasing returns to scale and evidence for passing price reductions to retailers more quickly than price increases.

In this article, I focus on explaining a seeming paradox: the coexistence of oligopolistic middlemen and gradually narrowing gross margin between farm-gate and UHT milk prices. McCorriston et al. (2001) offers a crucial insight to explain this seeming paradox. They offer theoretical support for the idea that if oligopolistic middlemen also enjoy increasing returns to scale, they will be able to tolerate a narrowing margin between main input—the farm-gate price—and output price—UHT milk price because increasing returns to scale enable them to sustain, or even increase, their net margins. I follow a two-pronged approach to test empirically whether their insight might be useful in explaining the seeming paradox in Turkey. First, I include the dairy industry labor productivity index on the right-hand side of cointegrating equations. Second, I include a dummy variable in the UHT milk price equation to account for the structural break in the data that also corresponds to two new competitors' entry into the fluid milk market. Cointegration tests for farm-gate and UHT milk prices alone indicate that they are not cointegrated. However, when the labor productivity index and dummy variables are included, the tests reveal a cointegrating relationship between these variables. As I discuss in Section 3.2, both UHT and farm-gate prices are both statistically and economically significant explanatory variables in the respective equations in the long-run relationship. The dummy variable that accounts for the structural break in the cointegrating relationship for UHT prices also significantly improves the fit of the long-term relationship equation. This finding is very significant because it suggests that appropriately constructed dummy variables (i.e., parallel to observed changes in industry structure), together with appropriate trend variables, can account for returns to scale. Moreover, these trend and dummy variables are easy to construct and utilize, so this empirical strategy to test for increasing returns to scale is easily replicable in other contexts.

Finally, these findings strongly suggest two arenas for further research. First, I show that even if the dairy processing firms are oligopolistic in nature, they behave competitively in the UHT milk market. However, competitive behavior on their part does not ensure a sufficiently steady supply of raw milk to meet demand. State Planning Organization (Devlet Planlama Teşkilatı, 2006) advocates for public regulation of the dairy market. Agricultural economists must be especially interested in exploring the effect on production of reducing farm-gate milk price volatility. Second, a complete picture of the welfare effects of oligopolies in the food processing and retailing sectors requires investigation of price transmission between dairy processing firms and retailers. A great deal can be learned by explicitly investigating price transmission between dairy processors and retailers.

APPENDIX

Enders and Siklos (2001) point out that Engle-Granger and Johansen cointegration tests and all of their variants are misspecified in the presence of asymmetry. Nevertheless, I follow Engle and Granger (1987) and estimate the classical error correction model (ECM) both for symmetric and asymmetric specifications in order to calculate short-run price elasticities. The crucial difference of the APT model (Equation A2) with an ordinary ECM (Equation A1) is that error correction terms (ECT) are split between positive and negative errors in the APT model. Unfortunately, in this specification it is possible only to test for the speed of adjustment because by assuming a cointegrating relationship in the long-run (Equations 2a and 2b), I implicitly assume that the magnitude of the difference between inflation-adjusted UHT milk price and farm-gate milk price is constant when controlling for labor productivity improvements and

TABLE A1. Error Correction Models With Alternative Specifications For Ultra-High Temperature (UHT) Milk

Dependent	d.UHT Symmetric	d.UHT Asymmetric	d.farm-gate Symmetric	d.farm-gate Asymmetric
Constant	−0.005	−0.007	0.001	0.005
D.farm	0.744***	0.740***		
D.UHT			0.102***	0.111***
D.prod	−0.002	−0.002	0.001	0.000
ECT_uht(-1)	−0.064			
ECT_uht_p(-1)		−0.046		
ECT_uht_n(-1)		−0.095		
ECT_farm(-1)			−0.153***	
ECT_farm_p(-1)				−0.236***
ECT_farm_n(-1)				−0.076
D.farm				
L1.	0.207	0.211	0.168*	0.166*
L2.	−0.087	−0.094	0.120	0.129
L3.	−0.434*	−0.434*	0.162*	0.170*
L4.	0.060	0.066	0.013	0.017
L5.	−0.005	0.001	0.119	0.131
L6.	−0.401*	−0.402*	0.196**	0.221**
L7.	−0.179	−0.177	0.007	0.030
L8.	−0.063	−0.067	0.128	0.140
D.UHT				
L1.	−0.044	−0.045	0.068**	0.074**
L2.	0.028	0.032	0.016	0.016
L3.	0.086	0.091	−0.072**	−0.071**
L4.	−0.035	−0.030	−0.051	−0.056*
L5.	0.009	0.013	0.005	0.000
L6.	−0.295***	−0.292***	0.051	0.044
L7.	0.028	0.030	−0.027	−0.033
L8.	0.037	0.038	−0.018	−0.031
D.prod				
L1.	0.002	0.002	−0.001	−0.001
L2.	0.000	0.000	−0.001	−0.001
L3.	−0.001	−0.002	0.002	0.002
L4.	−0.001	−0.001	−0.002	−0.002
L5.	0.005	0.005	0.000	0.000
L6.	−0.002	−0.002	0.001	0.001
L7.	0.001	0.001	−0.002	−0.002
L8.	−0.001	−0.001	0.002	0.001
Adj. R^2	0.1307	0.1245	0.2012	0.2098
B-P hett:	23.03***	23.92***	4.89**	4.65**
B-G LM	$0.085 \sim \chi(1)$	$0.135 \sim \chi(1)$	$0.937 \sim \chi(1)$	$1.437 \sim \chi(1)$
ARCH(LM)	$0.281 \sim \chi(1)$	$0.395 \sim \chi(1)$	$0.288 \sim \chi(1)$	$0.781 \sim \chi(1)$
D-W d-stat	2.0198	2.03	2.03	2.04
Durbin's alternative	$0.073 \sim \chi(1)$	$0.115 \sim \chi(1)$	$0.757 \sim \chi(1)$	$1.155 \sim \chi(1)$
Normality of residuals	15.48***	14.72***	18.43***	19.8***

structural breaks.

$$\Delta uht_t = \alpha + \sum_{j=1}^{8}\left(\beta_{1t}\Delta farm_{t-j+1}\right) + \sum_{j=1}^{8}\left(\beta_{2j}\Delta uht_{t-j+1}\right) + \sum_{j=1}^{8}\left(\beta_{3j}\Delta prod_{t-j+1}\right) + \varphi ECT_{t-1} + \varepsilon_t \tag{A1}$$

$$\Delta uht_t = \alpha + \sum_{j=1}^{8}\left(\beta_{1t}\Delta farm_{t-j+1}\right) + \sum_{j=1}^{8}\left(\beta_{2j}\Delta uht_{t-j+1}\right) + \sum_{j=1}^{8}\left(\beta_{3j}\Delta prod_{t-j+1}\right) + \varphi^{+}ECT_{t-1}^{+} + \varphi^{-}ECT_{t-1}^{-} + \gamma_t \tag{A2}$$

where Δuht_t is the change in inflation-adjusted wholesale UHT milk price, $\Delta farm_{t-j+1}$ is the change in inflation-adjusted farm-gate milk price, $\Delta prod_{t-j+1}$ is the labor productivity index, $ECT_{t-1}$is the lagged error correction terms from the long-term relationship, and ECT_{t-1}^{+} and $ECT_{t-1}^{-}$are positive and negative error terms from the long-term relationship.[9]

Table A1 shows that the coefficient estimates of contemporary change in farm-gate milk price (D.farm) are significant in both specifications. I chose the lag length as eight following the results of cointegration tests.[10] Heteroscedasticity is consistently detected in every specification. The *t* statistics in Table A1 are corrected for heteroscedasticity. For the symmetric (asymmetric) ECM for UHT milk price as dependent variable, 1 Turkish Lira (TL) increase in farm-gate milk price will lead to a 0.744 (0.740) TL increase in UHT milk price within a month and the short- run price elasticity at mean values is 0.213 (0.211). Short-run price elasticity estimates are almost twice that of long-run elasticity estimates, suggesting that initial reaction of UHT price to changes in farm-gate price overshoots the eventual reaction. After 3 months this overreaction dissipates. When the dependent variable is change in farm-gate prices, the coefficient estimates for concurrent change in UHT price are significant in both models. For the symmetric (asymmetric) ECM, 1 Turkish Lira (TL) increase in UHT milk price will lead to a 0.10 (0.11) TL increase in farm-gate milk price within a month and the short-run price elasticity at mean values is 0.357 (0.389). The short-run elasticity estimates are slightly less than long-run elasticity estimates suggesting that most of the adjustment in change in farm-gate prices in response to change in UHT prices takes place within a month.

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[9]I also report symmetric and asymmetric ECMs in Table A1 when the change in farm-gate prices is the dependent variable and change in UHT prices is the main independent variable.

[10]Cointegration tests are available from the author upon request.

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