FLUID MILK PROCESSORS MARKET POWER IN KOREAN DAIRY INDUSTRY AN APPLICATION OF THE AUTOREGRESSIVE DISTRIBUTED LAG APPROACH

By

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ABSTRACT

FLUID MILK PROCESSORS MARKET POWER IN KOREAN DAIRY INDUSTRY
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The Korean dairy market has become increasingly concentrated over the past several decades, being dominated by a few large dairy processors, which indicates that there is potential for the domestic milk processors to express market power. The objective of this study was to empirically measure the degree of oligopoly power of domestic white fluid milk processors by using the new empirical industrial organization (NEIO) approach which includes conjectural variation (CV). The autoregressive distributed lags (ARDL) bounds found that there exists a long-run equilibrium among the variables. The ARDL approach allows the estimation of the long-run and short-run coefficients using the ordinary least squares (OLS). The coefficients estimated in the long-run and short-run model are significantly different from zero, implying that the Korean milk industry is imperfectly competitive. Based on the oligopolistic power parameters there was an upward trend until 2003, then a downward trend, implying that the white fluid milk market has become more competitive over time. The Lerner's Indices calculated in the long and short-run are 14.9% and 3.7%, respectively, which indicate that the domestic white fluid milk processors have on average gained excess profits from 1985 to 2014 while exerting their oligopoly power.

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Chapter 1. Introduction

The Korean dairy industry rapidly developed up until 2002, increasing in the number of dairy cattle up to 553,000, but soon after the number of dairy cattle and farms continuously decreased. This was due to changes in government policies aimed at reducing the oversupply of raw milk. In spite of the decreasing trend, the improvement in dairy cattle genetics at the farm level made it possible to maintain raw milk production. Along with the changes at the farm level, the Korean milk processing market had also become mature. Some factors that contributed to this include: the lower national birthrate, the increase in breast feeding, and the growth of substitute beverage industries (Chang and Jeon, 2011)

Even though the Korean milk processing industry shows signs of stagnation, it is still one of the most dominant industries in Korea. In 2013, the market size of the milk processing and manufacturing industry, based on the value of shipments was 7,544 billion won ($6,618 million dollars) which accounted for 9.7% of the total food and beverage industry. The market size of fluid milk and other dairy products such as fermented milk, butter, cheese, and milk powder was 5,536 billion won ($4,856 million dollars) which accounted for 7.2% of the total food and beverage industry (KOSIS, Korea Statistical Information Service-The mining and Manufacturing Survey).

One of the characteristics of the domestic milk processing industry is that the concentration has increased. In general, if a firm is able to raise the market price of a good or service over its marginal cost and this firm can restrict other firms from entering, then there exists market power (Ahn, 2006). In addition, a high concentration ratio typically raises concern of firms abusing their market power. In Korea, the milk processing industry is considered to be oligopolistic because the CR3, the percentage of total sales made by the top three firms, has been between 60 to 80% since 2005 (Chang and Jeon, 2011). The domestic white fluid milk market is dominated by a few dairy processors: Seoul Milk Cooperative, Namyang Co., and Maeil Co. The market shares of these top three processors sum to 58% of the total domestic raw milk production (Jeon, 2011). Therefore, there is significant potential of market power exertion in the white fluid milk processor market.

Price increases in white fluid milk compared to price increases in raw milk imply that the domestic white fluid milk market has an imperfective market structure (Ahn, 2006). Since 1985, there have been ten shocks in the domestic raw milk price and white fluid milk price (the changes in raw milk price are determined by the government, Korea Dairy Committee, KDC). These two prices move together, indicating the changes in the raw milk price are directly transferred to the retail price, but the rate of increase is much higher than the raw milk price rate. For example, the raw milk price increased by 11.2% and 16.2% in 2004 and 2008, respectively, while the corresponding retail price increased by 21.8% and 35.1%. Though the retail price of the final white fluid milk is affected by production, marketing, and other inputs, the raw milk price is the most critical input that determines the retail price. As a result, these significant differences between the increasing ratios of the two prices may lead to excess profit for the milk processors.

Most empirical studies estimating market power in agricultural markets have used time series data such as retail prices of specific products, demand and supply, and other input prices. Empirical analysis based on time series assumes that these data are stationary, which means its mean, variance, and autovariance remain the same over time. However, it is well known that most business and economic time series data are not stationary. Gujarati (2003) shows that when nonstationary time series data are used, the results may produce a spurious regression. Therefore, testing whether the time series data are stationary or not is important.

Ahn (2006) and Jeon (2009, 2011) estimate the degree of oligopoly power of fluid milk processors in Korea based on the conjectural elasticity approach, where a conjectural elasticity indicates the degree of market power of milk processors. Although these papers used yearly or quarterly time series, tests for stationarity of the data or for long-run cointegration were not performed. To avoid spurious regressions this paper utilizes an autoregressive distributed lags (ARDL) model. Therefore, this study contributes to econometric robustness.

The main purpose of this study is to determine if Korea's domestic white fluid milk processors' market is imperfectly competitive. The three specific objectives of this research are to:

1. provide information about the current market structure of the white fluid milk market and background information on the Korean dairy industry;
2. empirically measure the degree of oligopoly power of domestic milk processors and determine the degree of market power changes as well as the level of margins of these milk processors changes during the research period;
3. and discuss the implications of the presence or not of market power in the processor industry.

Chapter 2 begins with background information about the Korean dairy industry, describing the supply and demand of raw milk and milk products, pricing system, market structure of dairy processing, and the manufacturing industry. It concludes with a literature review on measuring market power in the dairy industry. In Chapter 3, the ARDL bounds testing approach is presented along with the development of the empirical model. In Chapter 4, the data used for this paper is presented and empirical results are provided. Chapter 5 consists of conclusions and implications.

Chapter 2. The Korean Dairy Industry

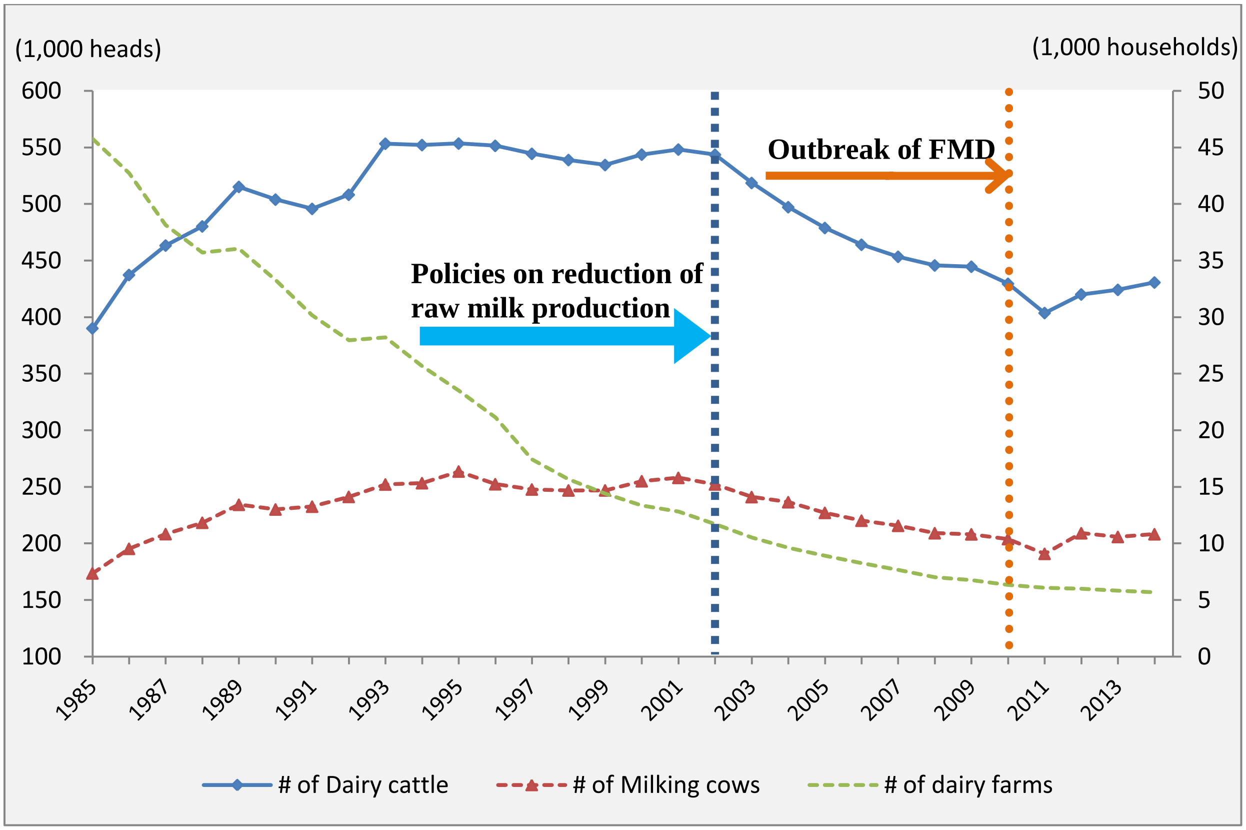
This chapter provides background information about the Korean dairy industry and reviews the literature on estimating market power in the dairy industry.

2.1. Supply and Demand of Milk and Milk Products

Figure 2.1 shows the trends of the number of dairy cattle and dairy farms in Korea. From 1985 to 2001 the number of dairy cattle and milking cows increased. The number of dairy cattle steadily increased by 40.5% until 2001, changing from 390,000 to 548,000 head. The number of milking cows increased by 48.6% during this same period, growing from 173,000 to 258,000 head. However, in 2002 the Korean dairy industry was faced with problems regarding the oversupply of raw milk so the government begun to adopt policies to reduce raw milk production. For example, in April 2002, the Korean Agricultural Ministry requested that each farm spontaneously slaughter 10% of their own dairy cattle. Moreover, in November 2002, a price differential system for surplus milk was introduced by the Korean Dairy Committee where the dairy farmers receive a lower price than the regular price for surplus milk. The government continued these policies for only one year. After that time, the number of cattle and milking cows consistently declined. At the end of 2010, there was an outbreak of foot-and-mouth disease (FMD) over the entire country which led to the culling of 36,397 dairy cattle. The Korean dairy industry experienced a temporary shortage of raw milk supply in 2011 because of this outbreak. Under these reduction policies and FMD outbreak, the number of dairy cattle and milking cows in 2014 were 431,678 and 208,205 head, respectively, a decreases of 21.4% and 19.3%, respectively compared to 2001.

Unlike the trends in the number of dairy cattle and milking cows, the number of dairy farms has been dramatically decreasing since 1985, changing from 45,760 farms in 1985 to 5,693 farms in 2014. The main reason for this downward trend is that many small-sized farms went out of the business due to the aging phenomenon in Korean farmers, the outbreaks of FMD, increased production costs caused by price increases of livestock feeds, and the increase in imports.

Figure 2.1. Change in the number of dairy cattle, milking cows, and dairy farms



*FMD: Foot-and-Mouth disease

Source: Dairy Statistics Yearbook (2014), Ministry for Food, Agriculture, Forestry and Fisheries & Korean Dairy Committee

The production and consumption of milk are shown in Table 2.1. The trend in domestic milk production is similar to the trend in the number of dairy cattle. Domestic milk production steadily increased by 2.5 times as much from 1985 to 2002, changing from 1,005 thousand tons to 2,536 thousand tons but thereafter, it decreased slightly until 2009 due to the policies for the reduction of raw milk production as mentioned above. The domestic milk production declined temporarily in 2011 because of the foot-and-mouth disease (FMD) at the end of 2010 but it returned to its former level in 2014. Then again, as the domestic agricultural market was opening, the imports of milk products continuously increased since 1995. The volume of imported milk products increased by about 8.6 times, changing from 195 thousand tons in 1995 to 1,683 thousand tons in 2014. The total production of milk steadily increased from 1,047 thousand tons in 1985 to 3,253 thousand tons in 2002, but afterward, it became stagnant until 2010. In 2014, the total production was 3,989 thousand tons which increased by 22.3% compared to that of 2010. This growth is largely due to increased imports and a slight decline in domestic production.

Total consumption and dairy products consumption per capita have a similar trend, along with the total production level. In 2014, the total consumption and per capita consumption of milk were 3,757 thousand tons and 72.4kg, respectively.

Table 2.1. Production and consumption of milk in South Korea

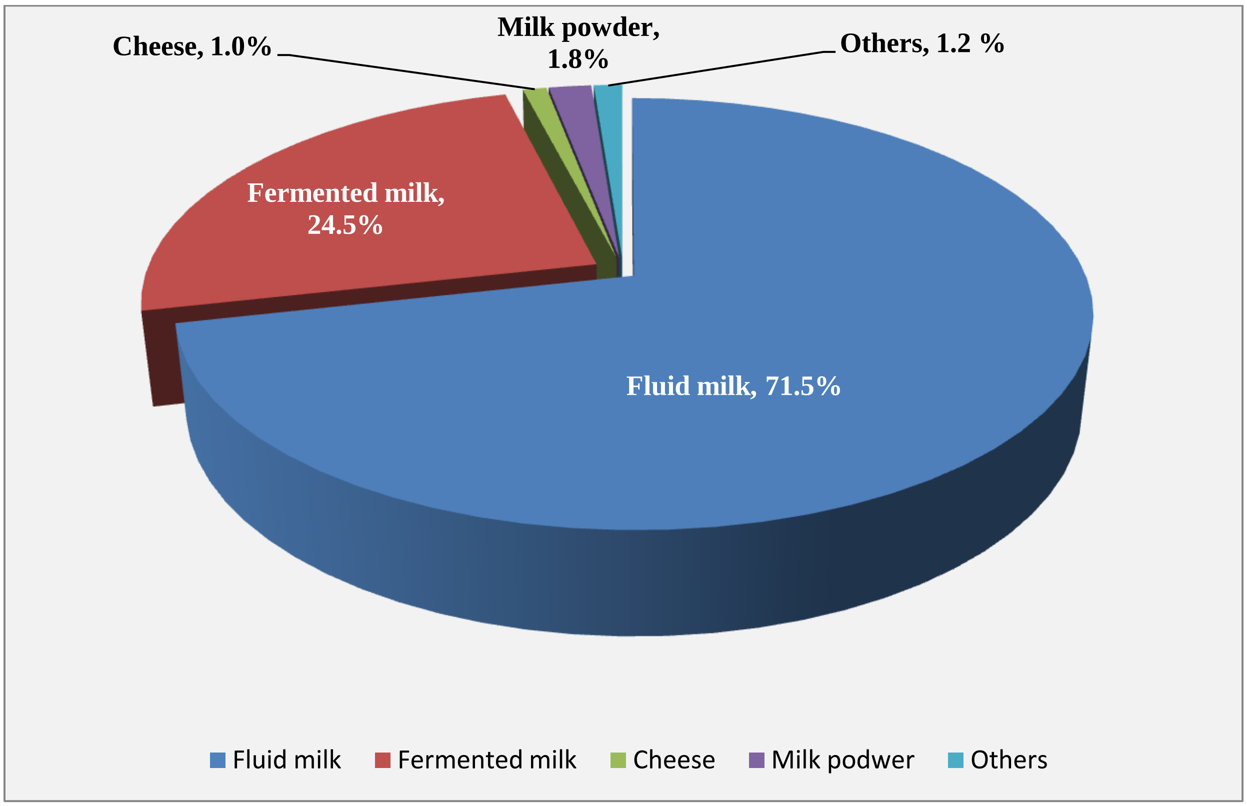
year	Beginning Stocks (1,000 ton)	Production (1,000 ton)	Imports (1,000 ton)	Total (1,000 ton)	Consumption (1,000 ton)	Ending Stocks (1,000 ton)	Per capita (kg)
1985	39.4	1,005.8	1.9	1,047.1	990.5	56.6	23.8
1990	150.3	1,751.8	0.0	1,902.1	1,879.0	23.0	43.8
1995	15.2	1,998.4	195.9	2,209.5	2,143.8	65.7	47.5
2000	43.6	2,252.8	639.6	2,936.0	2,811.5	124.5	59.6
2001	124.5	2,338.9	652.6	3,115.9	3,045.7	70.2	63.9
2002	70.2	2,536.6	646.5	3,253.3	3,092.3	161.0	64.3
2003	161.0	2,366.2	603.6	3,130.9	3,036.9	94.0	62.5
2004	94.0	2,255.5	842.1	3,191.5	3,123.5	68.0	64.0
2005	68.0	2,228.8	898.2	3,195.0	3,078.5	116.5	62.9
2006	116.5	2,176.3	882.3	3,175.2	3,121.7	53.5	63.5
2007	53.5	2,187.8	967.5	3,208.8	3,101.5	107.3	62.8
2008	107.3	2,138.8	885.1	3,131.2	3,034.9	96.3	60.9
2009	96.3	2,109.7	959.1	3,165.2	3,110.7	54.5	61.7
2010	54.5	2,072.7	1,134.8	3,262.0	3,249.4	12.7	64.2
2011	12.7	1,889.2	1,712.7	3,614.5	3,596.0	18.5	70.7
2012	18.5	2,110.7	1,414.4	3,543.6	3,451.8	91.7	67.2
2013	91.7	2,093.1	1,586.4	3,771.2	3,678.6	92.7	71.3
2014	92.7	2,214.0	1,682.8	3,989.5	3,757.0	232.6	72.4

Note: Milk production is based on the volume of milk which passed milk testing and imports are converted imported milk products that are turned into raw milk. Ending stocks are equivalent of milk powder left over from the previous year. Dairy products consumption per capita is calculated by dividing total consumption by total population.

Source: Dairy Statistics Yearbook (2014), Ministry for Food, Agriculture, Forestry and Fisheries & Korean Dairy Committee.

On average, approximately 70% of the raw milk produced is used for liquid milk products such as white fluid milk and flavored milk while the remainder is used for the manufacturing of fermented milk, cheese, butter, and milk powder (Dairy Statistics Yearbook, 2014). Figure 2.2 shows the proportion of each milk product to total domestic milk production and it indirectly indicates domestic raw milk utilization. About 71.5% (1,636 thousand tons) and 24.5% (572 thousand tons) of domestic raw milk produced in 2014 were utilized for fluid milk and fermented milk, respectively. That is, more than 90% of raw milk produced was used for fluid milk and fermented milk. While the proportion of cheese, milk powder, and others, is 1%, 1.5%, and 1.5%, respectively.

Figure 2.2. The proportion of each milk product to total milk production in 2014



Note: Others include cream, condensed milk and butter.
Source: Korea Dairy Committee (2014)

Table 2.2 shows the consumption of milk products over time. From 1985 to 2001 the consumption of white fluid milk increased by 2.3 times but after 2001, consumption stagnated. The consumption of flavored milk increased by 4.9 times from 1985 to 2004, but afterward, it steadily decreased, showing a change from 453 thousand tons in 2004 to 281 thousand tons in

2014. Along with these changes, the total fluid milk consumption increased to 1,829 thousand tons in 2003 but thereafter it declined to 1,637 thousand tons in 2014. The proportion of white fluid milk and flavored milk to total consumption in 2014 were 55.3% and 11.4%, respectively.

The fermented milk and other milk products have similar consumption patterns showing an upward trend until 2003 but then a downward trend until 2009. As the milk manufacturers actively participated in developing new products and public relations, the consumption of fermented milk and other milk products increased in 2014 by 28.7% and 51.6% compared to those of 2009, respectively (Dairy Statistics Yearbook, 2014).

Table 2.2. Consumption of milk products from 1985 to 2014

	Fluid milk			Fermented Milk (1,000 ton)	Other milk products (1,000 ton)
	White fluid milk (1,000 ton)	Flavored milk (1,000 ton)	Total (1,000 ton)		
1985	647.7	92.8	740.4	146.9	41.6
1990	1,242.1	94.3	1,336.5	352.9	76.7
1995	1,326.1	242.1	1,568.2	585.6	67.0
2000	1,447.4	224.1	1,671.5	529.2	133.0
2001	1,465.8	263.5	1,729.3	537.7	160.7
2002	1,362.1	302.2	1,664.3	540.4	160.4
2003	1,380.2	448.4	1,828.5	554.6	174.2
2004	1,328.3	452.9	1,781.2	524.6	161.8
2005	1,310.9	380.3	1,691.2	483.0	156.4
2006	1,343.7	339.9	1,683.6	504.4	158.2
2007	1,361.9	334.6	1,696.5	485.4	154.7
2008	1,351.5	350.8	1,702.3	454.7	152.3
2009	1,389.6	312.3	1,701.9	445.6	159.5
2010	1,362.0	279.2	1,641.1	502.1	182.2
2011	1,338.1	286.3	1,624.4	522.0	212.3
2012	1,405.1	280.2	1,685.3	557.7	206.0
2013	1,392.2	291.3	1,683.5	573.3	221.4
2014	1,356.3	280.7	1,637.0	573.4	241.9

Note: Fluid milk consumption is considered equal to the volume of its production because of the difficulties in long-term conservation of fluid milk. The consumption of fermented milk and others includes the volume exported. Cheese, cream, condensed milk, butter, and milk powder are included in other milk products category.

Source: Dairy Statistics Yearbook (2014), Ministry for Food, Agriculture, Forestry and Fisheries & Korean Dairy Committee.

2.2. Pricing System of Raw Milk and Price Changes of Milk Products

2.2.1. Pricing system of domestic raw milk

When the Uruguay Round of the World Trade Organization (WTO) agreement was signed in 1996, Korea formally opened their dairy market, but with minimum access quotas: relatively low within-quota tariff rates, and very high over-quota tariff rates. Before the WTO agreement was implemented, annual imports ranged between 4% and 9% of total domestic consumption. However, under the Uruguay Round (UR) agreement, total raw milk-equivalent imports increased sharply from 9% to 19% of consumption (Ahn et al., 2006).

Given the new market conditions, the Korean government passed "the Dairy Promotion Law" in 1997 and the Korean Dairy Committee (KDC) was established in 1999. Through the establishment of KDC, the government expected to control the supply and demand of raw milk as well as stabilize the prices of raw milk and milk products (Song et al., 2005). Since KDC was founded, the prices of raw milk have been determined by the KDC board based on production costs of raw milk[1] and other economic conditions. If the cost of production for raw milk increases by more than 5%, the KDC board will change the raw milk price. The milk processors can decide their own raw milk price but usually the raw milk price from the KDC is used.

As farm size in the dairy industry has increased and milk production per milking cow has improved, domestic dairy farms could be guaranteed relatively stable earnings with the higher raw milk price relative to the production costs. As shown in Table 2.3, there have been six increases of more than 5% in production costs since 2000. In 2004, a rise in the feed price led to an increase in the production cost of domestic raw milk by 5.3%. In addition, the production cost in 2008, 2011, and 2012 increased by 14.9 %, 12.0%, and 9.2% respectively. These sharp

[1] Production costs of raw milk based on 160 livestock farms are surveyed by Ministry for Food, Agriculture, Forestry and Fisheries and published every year, either in May or June.

increases were mainly driven by significantly greater world grain and crude oil prices (Chang and Jeon, 2011).

The constantly increasing cost of producing raw milk resulted in continuous increases in the raw milk price decided by KDC as well as the farm level price. The farm level price is always higher than production costs, regardless of the milk supply and demand. The milk producer price in 2011 was 894.6 won/ℓ ($0.85/ℓ equivalent), which increased by 44.0% in comparison to that of 2000. This price maintained the level of 600 won/ℓ from 2000 to 2004 and then the level of 700 won/ℓ until 2008. However, the recent price increases in the world grain and crude oil, and the exchange rate caused considerable rises in the production costs and then the farm level price increased by 5.7%, 9.9%, and 9.7% in 2008, 2009, and 2012, respectively. Prices paid to the milk producer include producers' incentives in addition to the basic raw milk price. These incentives are determined by the level of fat content, the number of germs, and somatic cells in the milk. If the fat content in milk exceeds 3.4%, producers can sell at a premium, while if it is lower than the standard level, then a penalty is assessed. Moreover, the farm level price varies depending upon the sanitation class which is determined by the number of germs and somatic cells in the milk.

Table 2.3. Production costs of raw milk and producers' price for milk

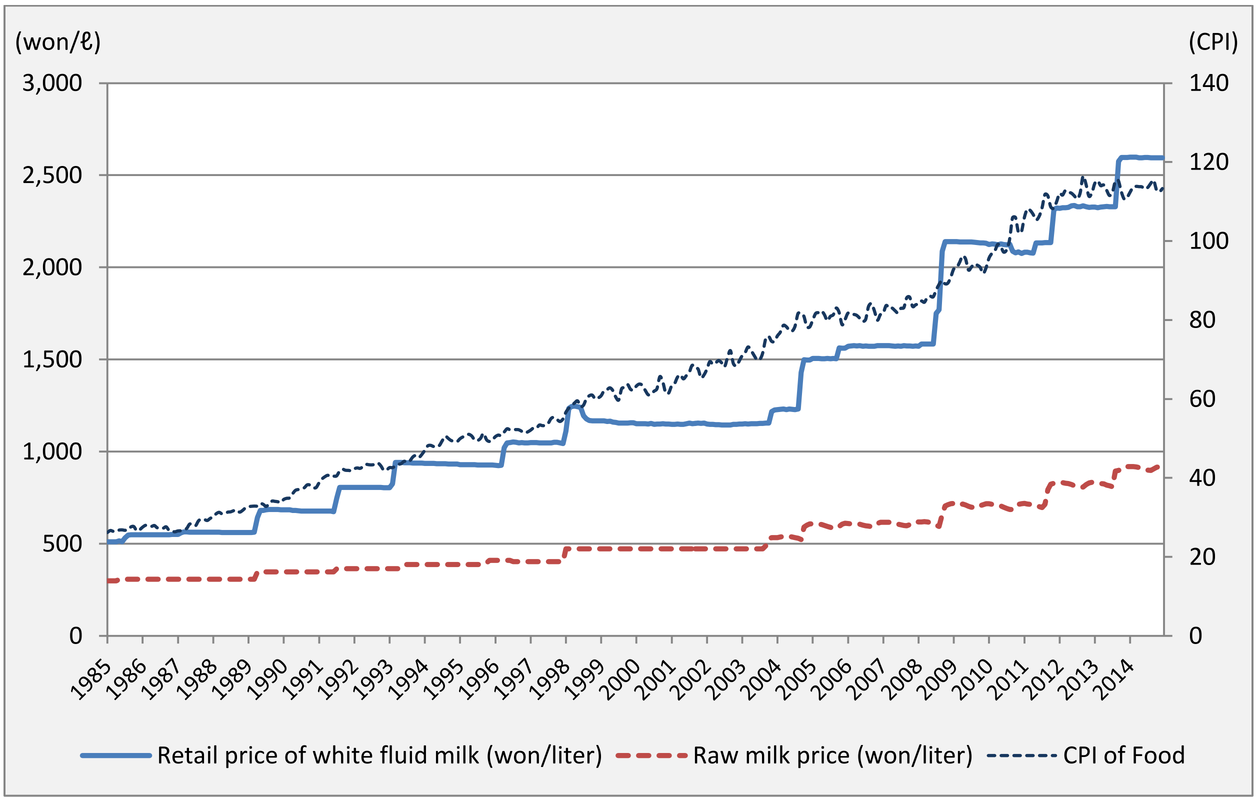
Year	Production cost (won/kg, ℓ)	Rate of change (%)	Producers' price (won/ℓ)	Rate of change (%)
2000	423	0.2	621.0	2.3
2001	446	5.4	628.9	1.3
2002	445	-0.2	633.5	0.7
2003	470	5.6	634.7	0.2
2004	495	5.3	662.2	4.3
2005	483	-2.4	716.8	8.2
2006	493	2.1	722.4	0.8
2007	509	3.2	728.5	0.8
2008	585	14.9	770.1	5.7
2009	614	5.0	846.6	9.9
2010	641	4.4	855.4	1.0
2011	718	12.0	894.6	4.6
2012	784	9.2	981.25	9.7
2013	807	2.9	1,022.16	4.2
2014	796	-1.4	1,088.09	6.5

Note: The unit of production cost was kg until 2002 and then it was changed to liter (1ℓ=1.03kg)
Source: Livestock Production Cost (2014) published by Statistics Korea and Korea Dairy Committee

2.2.2. Price changes of raw milk, white fluid milk, and total food

Figure 2.3 shows price changes in raw milk, white fluid milk, and total food using the Consumer Price Index (CPI) and Producer Price Index (PPI). The CPI for total food has steadily increased from 1985 to 2014 but the percentage change for each month is no more than 6%. The retail price of white fluid milk based on the CPI has a different trend over this time, showing periods of steep price hikes and stable prices repeatedly. The reason why the CPI for white fluid milk has a different moving pattern is that the retail price is primarily dependent on the changes in the raw milk price. The raw milk price is determined by the KDC as mentioned above and most dairy processors argue that the primary reason for a markup of the white fluid milk retail price is the increase in the raw milk price. In fact, the changes of the retail and raw milk prices seem to move together but their rates of change are considerably different. That is, white fluid milk retail and raw milk prices are steadily rising trends with retail price grows at a faster rate than the raw milk price[2].

Figure 2.3. Price changes of raw and white fluid milk, and total food



Note: Retail price of white fluid milk and raw milk price are based on the Consumer Price Index (CPI) and the Producer Price Index (PPI). These indices are transformed into prices, using the prices in 2010 as the base price.

[2] Over the period 1985q1-2014q4, the retail prices grew by 5.72% per annum, while the raw milk price increased by an annual average rate of 3.43%.

Table 2.4. clearly shows that the increasing rates of white fluid milk price are much larger than those of raw milk price except for the 2011 and 2013 rates. Over the period 1985-2014, there were ten shocks in the raw milk price and the corresponding increase of the retail price. In general, once the raw milk price increases, the milk processors tend to raise the retail price of white fluid milk consecutively for two or three time periods after the shock. For example, there was a 13.1% raw milk price shock in the second quarter of 1989. In contrast, the retail price of white fluid milk increased by 18.98% in the second quarter of 1989 and then increased again by 3.08% in the third quarter of 1989 compared to the prices of the previous quarters.

The relatively small increasing rate in 2011 may be a result of government regulation. The Korea Fair Trade Commission (FTC) reported in 2010 that the major milk processors in 2008 conferred on their selling prices of fluid milk and the FTC ordered them to take corrective action. The graphical analysis gives two implications: a linear trend may need to be included in the model and determining the lag length of the retail price changes is important.

Table 2.4. Rate of increase in white fluid milk price and raw milk price

	Time	Raw milk		White fluid milk	
		Price (won/ℓ)	Increasing rate (%)	Price (won/ℓ)	Increasing rate (%)
1	1/4 of 1985	298.64		510.00	
	3/4 of 1985	307.23	2.88	541.89	6.25
2	1/4 of 1989	307.23		561.38	
	3/4 of 1989	347.35	13.06	685.23	22.06
3	2/4 of 1991	347.49		676.28	
	4/4 of 1991	365.45	5.17	806.02	19.18
4	1/4 of 1993	365.45		857.11	
	3/4 of 1993	387.62	6.07	938.85	9.54
5	3/4 of 1995	387.48		928.02	
	3/4 of 1996	403.60	4.16	1,050.01	13.15
6	4/4 of 1997	403.60		1,047.22	
	2/4 of 1998	471.89	16.92	1,243.55	18.75
7	2/4 of 2004	539.19		1,229.32	
	4/4 of 2004	599.67	11.22	1,497.12	21.78
8	2/4 of 2008	612.20		1,584.09	
	4/4 of 2008	711.67	16.25	2,139.98	35.09
9	2/4 of 2011	705.97		2,113.11	
	1/4 of 2012	831.00	17.71	2,324.21	9.99
10	3/4 of 2013	818.75		2,410.55	
	4/4 of 2013	891.76	8.92	2,596.53	7.72

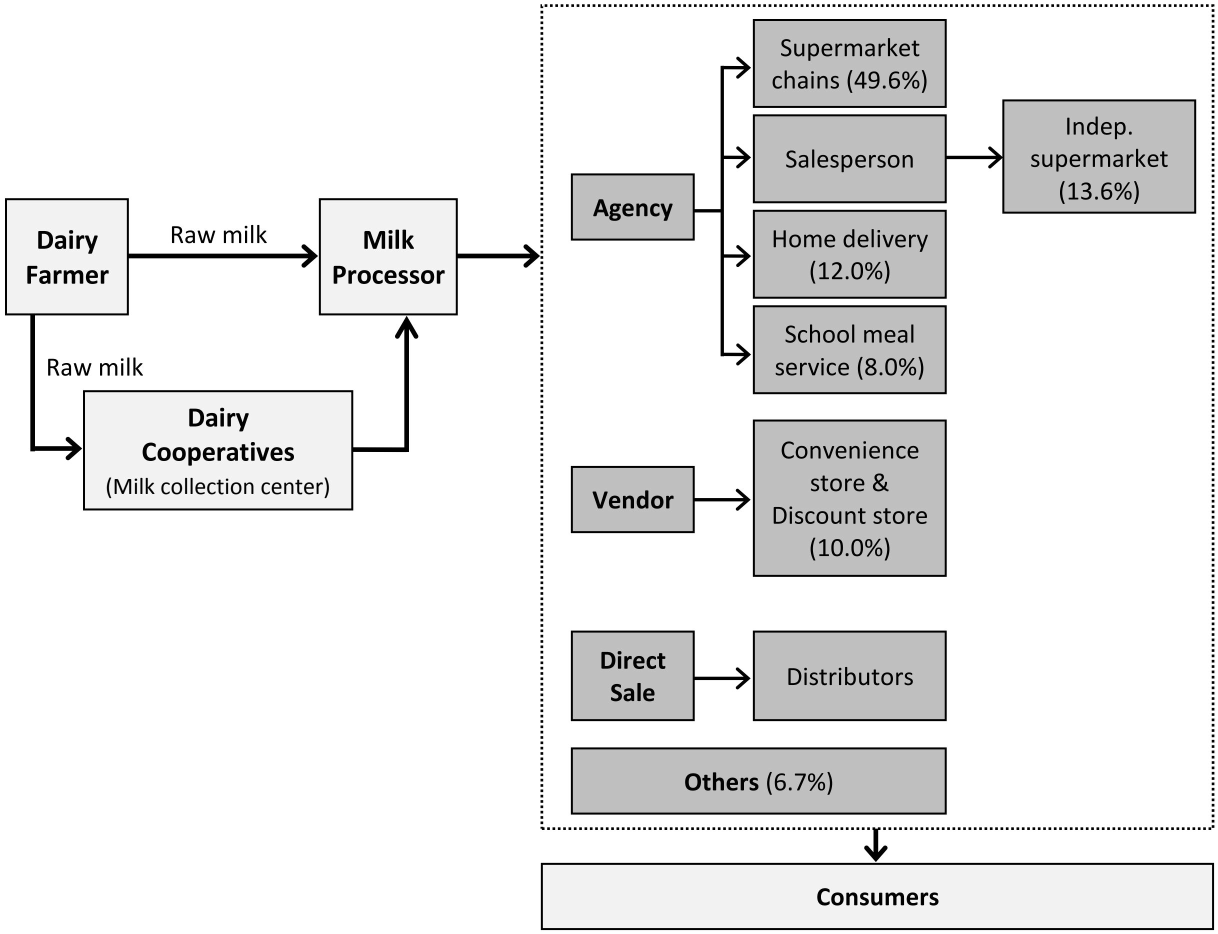
Note: Retail price of white fluid milk and raw milk price are based on the Consumer Price Index (CPI) and the Producer Price Index (PPI) and these indices are transformed into prices, using the prices in 2010 as the base price.

2.3. Distribution System of Raw Milk and Milk Products

The Korea Dairy Committee (KDC) attempted to unify the raw milk collecting systems but due to a low participation rate of dairy farmers and milk processors they failed. The proportion of the raw milk collected through the KDC was only 27% of the total raw milk in 2003 (Song et al., 2005). As a result, after the KDC was established, there existed two different collecting systems for raw milk, a direct collection by milk processors and indirect collection through the KDC. In the case of the indirect collection system, the processors were supplied the raw milk through the collection center assigned by the KDC.

Figure 2.4 shows that the processed milk products made by processors are consumed through a variety of distribution channels. The milk products are mainly distributed through milk agencies owned by milk processors, vendors, and direct sale channels (Yu, 2004). Agencies supply their milk products to supermarket chains (49.6%), to independent supermarkets through a salesperson (13.6%), to individual houses (12.0%), and to others including schools. In general, the top milk processors such as Seoul Milk Cooperative, Meail, and Namyang mainly tend to distribute milk products to supermarkets through their own agencies because their brand name is well known to the customers. Seoul Milk Cooperative also supplies the school meal service (Milk market report, 2013). On the other hand, the processors with lower brand awareness are apt to sell more to the home delivery market. The sales through vendors who supply convenience and discount stores account for 10% and the other channels include restaurants, coffee shops and bakeries (6.7%).

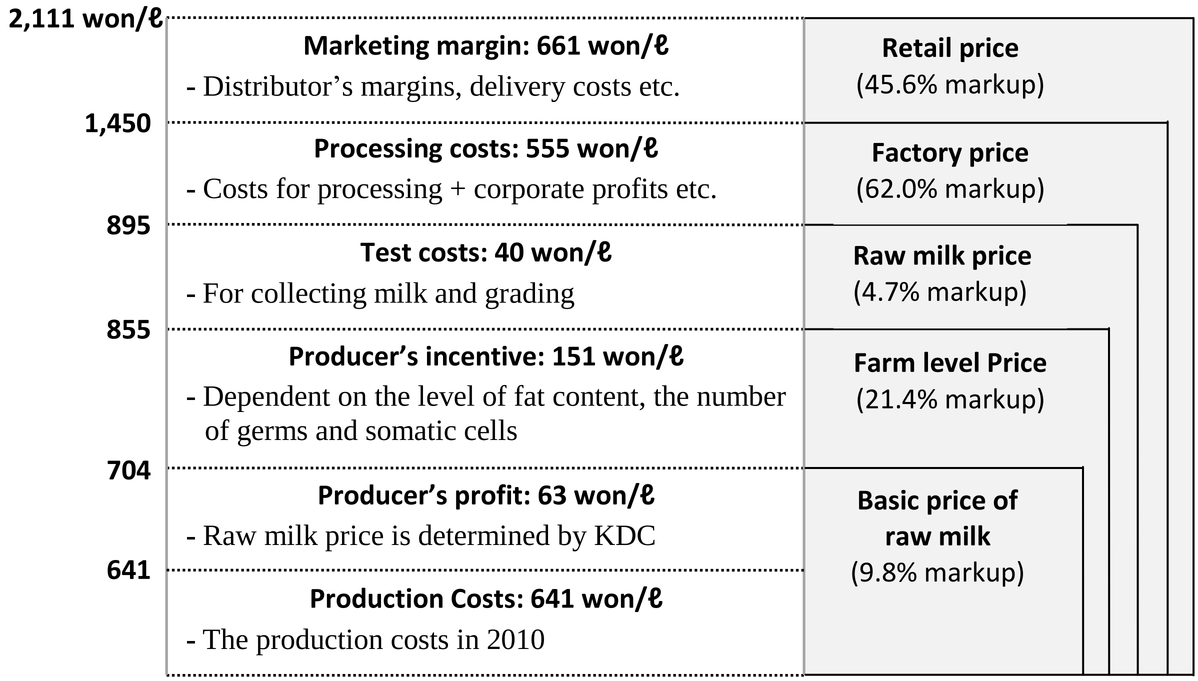
Figure 2.4. Distribution channel for raw milk and milk products



Note: The proportion of each distribution channel is based on white fluid milk.
Source: Milk market report (2013) published by the Ministry of Agriculture, Food and Rural Affairs and Korea Agro-Fisheries & Food Trade Corp., Fig. 4-1.

Figure 2.5 displays the marketing margin for each level of the fluid milk supply chain based on 2010 prices. For example, the production costs for raw milk was 641 won/ℓ and the basic raw milk price, determined by the Korea Dairy Committee, was 704 won/ℓ. The raw milk price accounted for 33.3% of the retail price of 2,111 won/ℓ. The price paid to the milk producer was 855 won/ℓ, which is reflective of the producer's incentive, which is determined by the level of fat content, the number of germs, and the somatic cells in the milk. The cost for collecting and grading milk was 40 won/ℓ. The factory price of white fluid milk was 1,450 won/ℓ which includes processing costs such as wages, packaging costs, and corporate profits. Therefore, the marketing margin for the distributors (wholesalers, retailers, supermarkets, and department stores) was 661 won/ℓ which accounts for 31.3% of the retail price of 2,111 won/ℓ.

Figure 2.5. Pricing decision system of white fluid milk, 2010



Source: Chang, J. and Jeon. (2011) "Mid to Long-Term Development Strategy for Milk Processing Industry", Fig. 2-5.

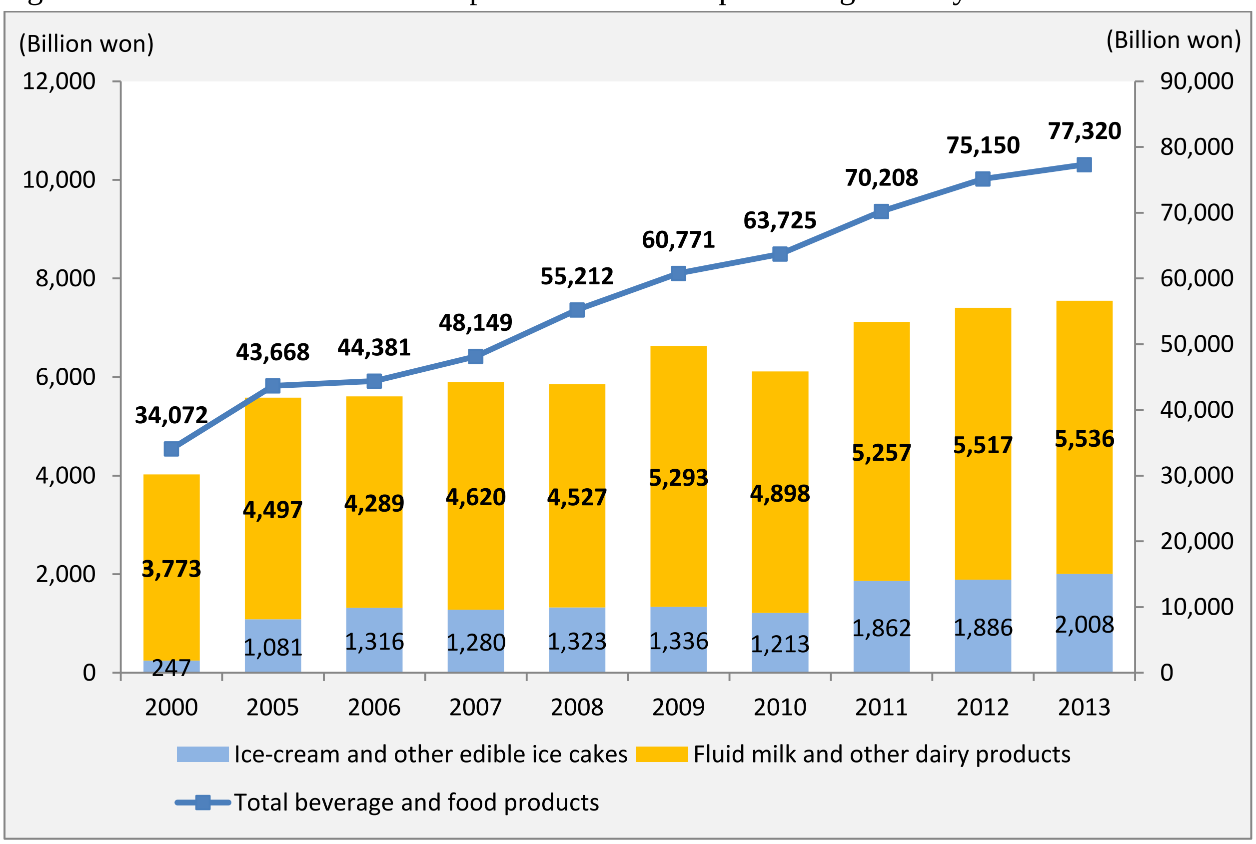
2.4. The Korean Milk Processing Industry

2.4.1. Market size of milk processing and manufacturing industry

The Korean domestic milk processing and manufacturing industry can be classified into two groups, manufacturers of fluid milk and other dairy products and manufacturers of ice cream and other edible ice cakes. As noted in Figure 2.6, the shipment value for the milk processing and manufacturing industry was 7,544 billion won in 2013 (9.7% of the total food and beverage products) and it has continuously grown since 2000 at an annual growth rate of 5.9%. The shipment value of fluid milk and other dairy products went up to 46.7% and that of ice cream and other edible ice cakes increased by 8.1 times as much compared to the 2000 sales.

The value of shipments of fluid milk and other dairy products in 2005 increased by 19.2% as compared with that in 2000 but thereafter it seems to have stagnated until 2008. This stagnation was caused by decreases in fluid milk and milk powder consumption from lower birth rates and increased breast-feeding, as well as the development of substitute beverage industries. However, the shipment values had an upward trend since 2008 and may be attributed to the increases in milk product retail prices in 2008 and 2011 as well as increases in milk consumption.

Figure 2.6. Trend in the value of shipments of the milk processing industry



Source: KOSIS, Korea Statistical Information Service-The mining and Manufacturing Survey

2.4.2. Market concentration of milk processing industry

In general, a firm has market power if it is able to raise the market price of a good or service over its marginal cost and this firm can restrict market entry of other firms. To evaluate the degree of concentration and to determine if firms are expressing market power, economists use several different methods. One of these methods, the four-firm concentration ratio (CR4), measures the total market share of the top four firms in an industry[3].

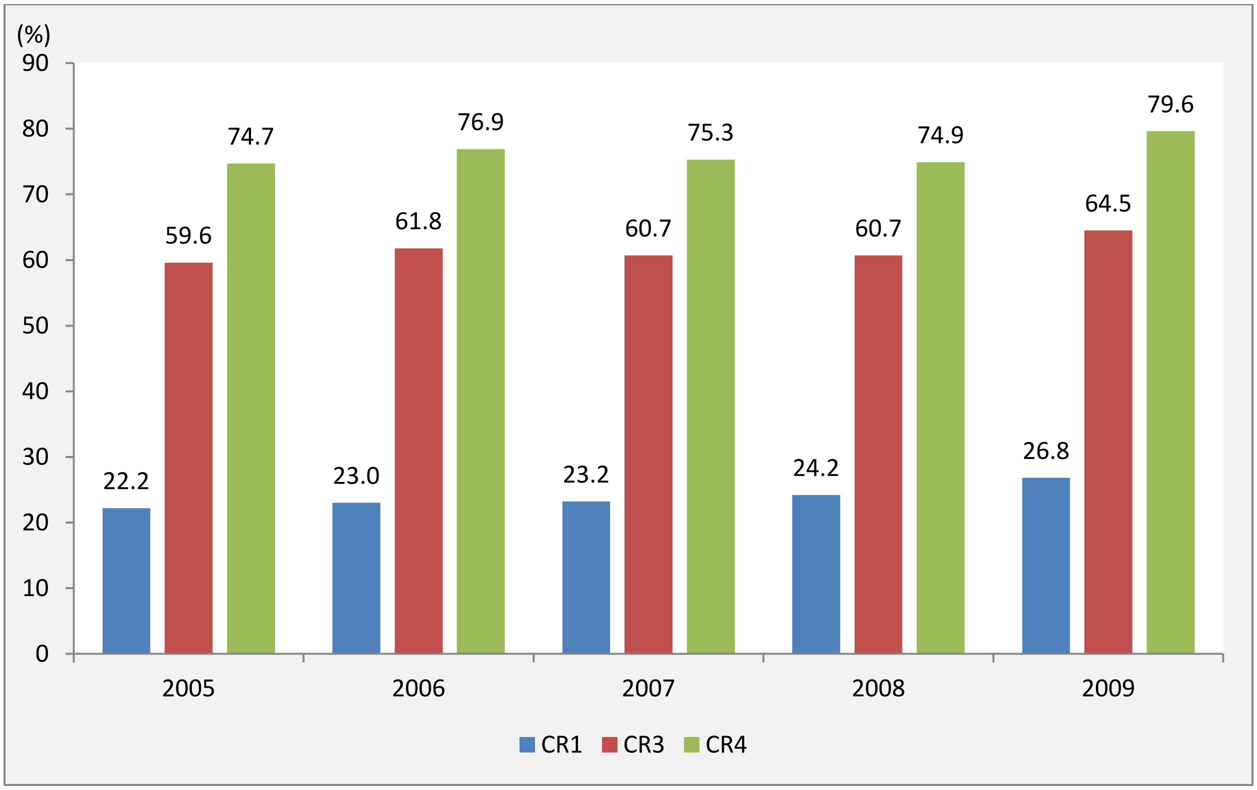
In 2009, the concentration ratio of the top milk firm was 26.8% and the CR4 accounted for almost 80%. The top four firms are Seoul Milk Cooperative, Korea Yakult Co., Namyang Co., and Maeil Co. The concentration ratio of the top milk processor, Seoul Milk Cooperative, increased from 22.2% in 2005 to 26.8% in 2009. In addition, the ratio of the top four firms also increased by 3.9%, changing from 74.7% in 2005 to 79.6% in 2009. Figure 2.7 shows that the degree of market concentration in the domestic milk processing industry is gradually growing.

This high concentration ratio seems to come from inherent characteristics of the domestic milk processing industry. The processed milk products have a relatively short shelf life and the technological level of Korea for making high value milk products, such as cheese and butter, is lower compared to other developed dairy countries. For these reasons, Korea is restricted from exporting milk products abroad and thus, the Korean milk processing industry has to be domestically oriented. In addition, new firms are likely to face a high barrier of entry into the milk processing industry due to the large initial investment. Although the Korean milk industry is saturated, domestic milk processors fiercely compete with each other to survive, actively

[3] In Korea, an industry with a concentration ratio of 80% to 100% is considered to be a oligopolistic market and a concentration ratio of 60% to 80% indicates an monopolistic competitive market. If concentration ratio is less than 60%, the industry is considered to be a competitive market (Chang, J. and Jeon, 2011). Song et al., (2005) reported the CR4 for the milk processing industry in 2003 and 2004 was 80%, meaning it is an oligopolistic market structure approaching monopolistic.

promoting and developing new milk products. At the same time, they also have to compete with the relatively cheap imported milk products.

Figure 2.7. Market concentration ratio of the milk processing industry



Source: Chang, J. and Jeon, (2011) "Mid to Long-Term Development Strategy for Milk Processing Industry", Fig. 3-2.

2.5. A Brief Summary of Studies on Estimating Market Power in the Dairy Industry

While empirical studies of market power in agricultural markets are common, there are relatively few studies that estimate the dairy industry market power and even fewer studies that have considered milk processors, specifically.

2.5.1. NEIO and Conjectural Variation (CV) models

Ahn (2006) estimated the market power of the Korean milk processors using the New Empirical Industrial Organization (NEIO) method that includes conjectural variation (CV). He showed that the retail price of fluid milk had excessively increased compared to the raw milk price. For example, the retail price of fluid milk in 2004 increased by 19% while the increase in raw milk price was only 13%. The computed coefficients of the market power parameter were significantly different from zero and were dependent on the price elasticity of the fluid milk demand. He concluded that the Korean fluid milk industry is far from competitive market and the domestic milk processors had on average gained more than 15% excess profits in the years from 1975 to 2004.

Jeon (2009) estimated the degree of oligopoly power of fluid milk processors in Korea also using the NEIO approach. He measured the degree of oligopoly power in two ways: first by directly calculating the Lerner's Index from the estimated value of marginal costs and second by indirectly recovering the degree of oligopoly with information from the demand side. The average value of the directly calculated Lerner's Index for white fluid milk was 0.11. With the indirect way, he obtained 0.15 for the oligopoly power of white fluid milk processors. This supports that the market structure of white fluid milk is more likely to be imperfectly competitive. However, this paper also presented the idea that the estimate obtained by the NEIO could be biased, citing Corts (1999) paper. Therefore, the market power parameters from the NEIO approach could be altered if there is a positive demand shock and there is a possibility of change in the conjectural variation parameters from seasonal demand shocks. The findings show that the slopes of the supply relation in the first and fourth quarter are relatively steeper, while the slopes in the second and third quarter are flatter, meaning the domestic fluid milk market becomes more competitive in summer than other seasons, however the oligopoly structure still persists. Also, he calculated the difference in the conjectural elasticity and it was 0.06, which is quite large compared to the Lerner's Index for white fluid milk (0.11). Therefore, the NEIO estimate for white fluid milk is underestimated. However, this paper did not discuss why the milk market structure becomes more competitive in summer season.

There are several studies outside of Korea that inform the model development for this paper. Chidmi et al. (2005) stated that private market power in the processing and retailing sectors exist together with the public policies that influence milk prices. This paper evaluated retail oligopolistic market power and the effects of the Northeast Dairy Compact (NEDC) in the Boston area, where the NEDC has the authority to regulate the farm price for fluid milk. This paper measured the retail market power using a market conduct parameter and the Lerner Index. They also examined if the conduct parameter changed depending on the existence of the NEDC, using the four scenarios: compact with and without private market power and no compact with and without private market power. The empirical results show that the conjectural variation elasticity calculated in both pre-and post-NEDC cases was significantly different from zero at the 5% level, implying the fluid milk market in Boston was not perfectly competitive. They also found that the NEDC dummy variable was statistically significant and the price increase caused by oligopoly power was seven times higher than those caused by the NEDC. As a result, approximately 25% of the retail price (about $0.75/gallon) was due to the impact caused by retail oligopoly power.

Cakir et al. (2011) found that concentration in milk marketing, processing, and retailing in the United States raises the potential for firms in the milk supply chain to exert market power. This paper adopts an econometric model of a vertical relationship between cooperatives and processor-retailers to measure the degree of market power. The model allows both diary co-operatives and processors to potentially exercise market power in their respective output markets, but processors are assumed to be price-takers because the Federal Milk Marketing Order

(FMMO) regulations making processors price-takers in the market for their primary input. Also, the authors derive the NEIO model that is linear in markups and utilize simultaneous-equations model (SEM). The key finding was that the conduct parameter for dairy cooperatives was small, but the markup level was fairly large, 9%, because the derived demand for milk facing cooperatives is very inelastic. This indicates that cooperatives are able to exert their market power to increase the price of milk purchased by fluid milk plants by approximately 9% over the minimum price set by the FMMO regulations. On the other hand, the retail demand for fluid milk is also quite inelastic, since the estimated conduct parameter for processor-retailer is relatively small the retail markup level is less than 1%.

Sckokai et al. (2009) estimated the market power of retailers for the supply chains of Italian hard cheeses, Parmigiano Reggiano (PR), and Grana Padano (GP) which are made from raw milk. This study assumes that retailers who buy the grana cheese from ripeners and then sell PR and GP to final consumers exercise market power in both the downstream and upstream market. In order to measure the degree of market power exerted by the retailers, the authors jointly estimated the market power parameters with supply and demand elasticities using the Generalized Method of Moments and monthly time series from 2002 to 2006. One of the key findings was that the estimated degree of oligopoly power was stronger for PR than for GP (0.87 vs 0.52) and this provides evidence of the existence of downstream market power by retailers over final consumers. This study concluded that one of the reasons that PR has a higher oligopoly power than GP was due to the fact that Italian consumers perceives PR as the highest quality grana cheese leading to higher retail price premiums. On the other hand, this study fails to find evidence of upstream market power (over the processors and ripeners).

Suzuki et al. (1993) reported that in Japan, like many countries, there was price discrimination between fluid and manufacturing milk since the fluid milk demand was more inelastic than the manufactured product demand. In addition, they show that the market power exerted by prefectural milk marketing boards, which almost have a monopoly on the milk supply, strongly influences prices that producers receive and that buyers pay. This study estimated the degree of competition in the Japanese milk market using conjectural variations (CV) and measures the effect of a milk support price reduction. The calculated values of the conjectural variations (CV) indicated that one region was more competitive than the other and both regions are steadily decreasing over time indicating increasing competitiveness. Under the assumption of one-time support price reduction, this study concludes that the fluid milk price will decrease less under the current imperfectly competitive structure than it would under a competitive structure.

2.5.2. Other models

Jeon (2011) measured the degree of oligopoly power of domestic milk processors in Korea using natural experiments by examining the impacts of an exogenous policy shock. In this paper, the raw milk price was considered to be an exogenous shock to the retail white fluid milk price because it was decided by the government. He estimated the net impact of a raw milk price shock on the retail price of fluid milk, using a difference-in-difference method. This method controls for all the other factors which affect retail price of fluid milk, such as wage and gas expenditure, while allowing the raw milk prices to vary. To do this, the sample was divided into two time series groups: one is a treatment group, which was influenced by the raw milk price shock, and the other was a control group, which was unaffected by the shock. The average of the

Lerner Indices calculated from eight shocks in raw milk prices was around 0.7, excluding 2003. He also considered a dynamic game to check the robustness of the estimates obtained from the static game. Given the behavior of the marketers, the results were not significantly different from those in a static game.

Suzuki el al. (1994) developed an imperfect competition model with an endogenous fluid price differential to evaluate the market impacts of U.S. dairy industry deregulation. This study shows that the Class I (fluid) differential actually cannot be exogenous because there exists over-order payments resulting from negotiations. To measure the degree of imperfection in the U.S. milk market, they estimated the effective fluid price which was expressed as the sum of the M-W price (the average price paid for manufacturing grade (Grade B) milk by manufacturing plants in Minnesota and Wisconsin), the minimum fluid price differential, and any over-order payment. The result from this study shows that the estimated parameter (θ), an aggregate indicator of the degree of competition in the U.S. milk market, has been declining over time, changing from 0.077 in 1977 to 0.055 in 1990. This estimation indicates that the U.S. milk market has become more competitive over time, leading to two interpretations. First, competitive pressures in the fluid milk market have increased over time as transportation technology has improved and reserves of milk in areas other than Minnesota and Wisconsin have increased. Second, dairy farmers have tried to reduce the competitive pressures by merging cooperatives and milk marketing orders while the size of the manufacturing plants which produce non-fluid dairy products have become larger.

Prasertsri and Kilmer (2008) stated that as a result of economies of size, food processors generally had a bargaining advantage over independent farmers, but marketing cooperatives were established to even out the bargaining position of independent farmers. This study estimated the relative bargaining power of one milk marketing cooperative (Southeast Dairy Cooperative, Inc. (SDC)) and several fluid milk processors in Florida from 1998 and 2004, using a Nash bargaining model to analyze the negotiated price in the fluid milk market which acts like a bilateral monopoly. They found that the monthly bargaining strength of SDC was higher than that of the processors for all twelve months. This study explained that one of the reasons why the SDC has the majority of the bargaining power was that processors are more impatient than the SDC because the processors have buyers who need a continuous supply of dairy products. In addition, all of the milk for the processors was provided by the SDC through a full supply contract and the milk marketing cooperative has higher bargaining power because it has the ability to sell milk to other markets. Even though Florida fluid milk processors could purchase milk from other areas, they would have to pay additional for the transaction costs of importing the milk into Florida.

Ahn (2006) states that the federal milk marketing orders, which were established by the Agricultural Marketing Agreement Act of 1937, have been motivated in part to prevent milk processors from exerting their potential market power on raw milk markets and this study identified the interaction between market power and the price regulations of the milk marketing orders. This study provides two ways to assess the political market power of milk producers: to assess the political market power by comparing the announced price differentials to the optimal ones that give maximum profits to producers; and, to estimate the political market power by deriving the welfare weights for milk producers using the policy preference functions. The results suggest that milk producers have more political power than buyers, but their political power was relatively small compared to monopoly power in setting prices. In addition, this study provided evidence of the existence of bilateral market power between raw milk buyers (milk bottlers) and sellers (dairy cooperatives) by using a fixed-effects fluid milk pricing equation to investigate responses to minimum prices in milk marketing orders. The key result indicates that minimum price policy has contributed to the welfare of milk producers by raising the bargained prices. Finally, the author measured the relative bargaining power of buyers by using the ratio of the difference between the observed bargained price and regulated minimum price to the difference between the upper bound of price for bargaining and regulated minimum price. The estimated results imply that the relative bargaining power of dairy cooperatives in regional raw milk markets was significant, but small compared to the bargaining power of milk bottlers.

Cotterill and Samson (2002) estimated a brand-level demand system for American cheese products and evaluate unilateral and coordinated market power for cheese brands using quarterly market level data of 33 U.S. cities from 1988 to 1992. To measure the market power of the cheese brands, they used the Cotterill Index (CI), which measures the total market power of a brand, and the Chamberlin Quotient (CQ), which measures the proportion of market power caused by tacit collusion. The American cheese products can be represented by five brands: Kraft, Velveeta, Borden, Private label, and “other brands”. The first two brands were owned by the Philip Morris/Kraft Company. Based on the price reaction elasticity, this study strongly argues that the Philip Morris Company was the price leader in the American cheese category. In addition, the calculated Cotterill Index (CI) shows that Kraft was the most powerful brand, exerting 61% of its potential market power. The Chamberlin Quotient (CQ) for Kraft showed that Kraft received 22.6% of its power from its leadership position. On the contrary, Private label gains only 0.8% from cooperation in the category pricing game.

Chapter 3. Unit Roots and Cointegration Tests, and Autoregressive Distributed Lag (ARDL) Bounds Testing Approach

Most of the studies for the estimation of the market power used time series data such as market prices, demand and supply, and other input prices, where the empirical analysis assumes that the underlying time series is stationary, meaning its mean, variance, and autocovariance remain constant over time. Even though most business and economic time series data are not stationary these previous studies overlooked a potential spurious regression problem caused by using nonstationary time series data. That is, if nonstationary time series data is used in a regression model then the coefficient of determination (R^2) is likely to be high, even though the series are actually independent of each other (Gujarati, 2003). Therefore, this paper included stationarity tests for the time series data using three different unit root tests which were the Augmented Dickey-Fuller (ADF), Philips-Perron (PP), and the DF-GLS test. In addition, cointegration analysis was considered as a pre-test to avoid spurious regressions.

3.1. Testing for Unit Roots and Cointegration

3.1.1 The Augmented Dickey-Fuller test (ADF test)

In 1979, Dickey and Fuller developed a test of stationarity using the fact that testing for stationarity is equal to checking whether a unit root is present in an autoregressive model. This test, Dickey-Fuller (DF) test, assumes the error terms u_t are uncorrelated. However, if the error terms u_t are correlated, they developed an augmented version of the Dickey-Fuller (DF) test, known as the augmented Dickey-Fuller (ADF) test (Gujarati, 2003). The ADF test follows the same procedure as the DF test but it adds the lagged values of the dependent variables. The basic idea of the augmented version was to include enough lagged difference terms so that the error term was serially uncorrelated. Here, an appropriate lag length for the augmented regression can be decided by using specific criteria such as the Akaike Information Criterion (AIC) and the Schwarz's Bayesian Information Criterion (SBIC). The ADF test was estimated in three different forms which are a pure random walk, a random walk with drift and a random walk with drift around a stochastic trend.

Model (i): A pure random walk $$\Delta Y_t = \delta Y_{t-1} + \sum_{i=1}^{p} \alpha_i \Delta Y_{t-i} + u_t \tag{1.1}$$

Model (ii): A random walk with drift $$\Delta Y_t = \beta_1 + \delta Y_{t-1} + \sum_{i=1}^{p} \alpha_i \Delta Y_{t-i} + u_t \tag{1.2}$$

Model (iii): A random walk with drift around a stochastic trend:

$$\Delta Y_t = \beta_1 + \beta_2 t + \delta Y_{t-1} + \sum_{i=1}^{p} \alpha_i \Delta Y_{t-i} + u_t \tag{1.3}$$

where Δ is the first-difference operator, β_1 is a drift parameter, p represents the lag structure, t is time or trend variable and u_t is white noise error term. In each case, the null hypothesis is that $\delta = 0$; that is, there is a unit root, meaning the time series data is nonstationary. The alternative hypothesis is that δ is less than zero; that is, the time series data is stationary. If the calculated absolute value of the τ statistic from each case is higher than the DF or Mackinnon critical τ values, we can reject the hypothesis that $\delta = 0$, indicating the time series data is stationary.

3.1.2. The Philips Perron (PP) test

The DF test assumes that the error terms u_t are independently and identically distributed. Under the assumption of correlation in the error terms, the ADF test adjusts the DF test to deal with the possible serial correlation problem by adding the lagged difference terms of the dependent variable (Gujarati, 2003). Philips and Perron (1988) offered an alternative method for correcting the serial correlation in the error terms without adding the lagged difference terms. Basically, the PP test uses nonparametric transformations of the τ statistics from the original DF or ADF test and the transformed statistics, the z statistics, have DF or ADF distributions (Phillips and Perron, 1988). The PP test uses three different models depending on if there is a constant and a time trend.

Model (i): Without a constant and trend $$Y_t = \delta Y_{t-1} + u_t \tag{2.1}$$

Model (ii): With a constant and without trend $$Y_t = \beta_1 + \delta Y_{t-1} + u_t \tag{2.2}$$

Model (iii): With a constant and trend $$Y_t = \beta_1 + \beta_2(t - \tfrac{1}{2}T) + \delta Y_{t-1} + u_t \tag{2.3}$$

where T is the sample size.

3.1.3. The DF-GLS test

Eliott, Rothenberg and Stock (1996) proposed the modified Dickey-Fuller test statistic for a unit root in which the time series data has been transformed by a generalized least squares (GLS) regression and it is known as the DF-GLS test. This test has substantially improved power when an unknown mean or trend is present (Elliott et al., 1996). There are two possible alternative hypotheses: y_t is stationary around a linear trend or y_t is stationary with no linear time trend. To perform the DF-GLS test under the first alternative hypothesis, the DF-GLS estimation is performed by generating the new variables: $\tilde{y}_t$, x_t, and z_t, where,

$$\tilde{y}_t = y_t - \alpha^* y_{t-1}, \qquad t = 2, \dots, T$$

$$x_t = 1 - \alpha^*, \qquad t = 2, \dots, T$$

$$z_t = t - \alpha^*(t-1)$$

and $\alpha^* = 1 - (13.5/T)$, $\tilde{y}_1 = y_1$, and $x_1 = z_1 = 1$ An OLS regression is then estimated for the equation,

$$\tilde{y}_t = \delta_0 x_t + \delta_1 z_t + \epsilon_t$$

The OLS estimators $\hat{\delta}_0$ and $\hat{\delta}_1$ are then used to remove the trend from y_t; that is, generate

$$y^{*} = y_t - (\hat{\delta}_0 + \hat{\delta}_1 t) \tag{3.1}$$

The last step is to perform an augmented Dickey-Fuller test on the transformed variable by fitting the OLS regression,

$$\Delta y_t^{*} = \alpha + \beta y_{t-1}^{*} + \sum_{j=1}^{k} \varsigma_j \Delta y_{t-j}^{*} + \epsilon_t \tag{3.2}$$

and then test the null hypothesis $\mathrm{H}_0\colon \beta = 0$ by using tabulated critical values.

Under the second alternative hypothesis, it follows the same procedure except for defining $\alpha^{*} = 1 - (7/\mathrm{T})$, eliminating *z* from the GLS regression and computing $y^{*} = y_t - \hat{\delta}_0$. In this study, the test is performed only for the first alternative hypothesis assuming the time series data used for the empirical analysis has upward or downward trends.

3.1.4. The Engle-Granger (EG) and the Augmented Engle-Granger (AEG) test

As mentioned above, the test on whether the time series data is stationary or not is important because the regression analysis using nonstationary time series data can lead to false results. However, even though two time series are individually *I*(1) (integrated of order 1), their linear combination can be stationary by the cancelling out of the stochastic trends between the two series. If this is the case, then the two variables are cointegrated, implying the regression between the two variables would be meaningful, not spurious. Economically speaking, two variables will be cointegrated if they have a long-term, or equilibrium, relationship between them (Gujarati, 2003). This study used two different methods for the general cointegration test: the Engle-Granger (EG) test, the augmented Engle-Granger (AEG) test, where EG and AEG perform tests on the residuals of the DF and ADF unit root test.

3.2. Autoregressive Distributed Lag (ARDL) Model and Test

The ARDL bounds testing approach initially proposed by Pesaran and Shin (1999), and Pesaran et al. (2001) has several advantages over the traditional cointegration approaches, such as the Engel-Granger (1987) two-step residual-based procedure and the system-based reduced rank regression approach pioneered by Johansen (1988, 1995) (Majid and Hafasnudin, 2010). One of the advantages is that the ARDL approach can be applied irrespective of whether the time series are purely $I(0)$, purely $I(1)$ or fractionally integrated (Pesaran and Pesaran, 1997; and Bahmani-Oskooee and Ng, 2002). This circumvents the problems arising from nonstationary time series data (Laurenceson and Chai, 2003). Another advantage for using the ARDL approach is that it incorporates sufficient lag-length into its analysis to capture the data generating process in the case of general-to-specific modeling framework (Laurenceson and Chai, 2003). Finally, the ARDL approach can be applied to studies that have a small sample size (Narayan, 2005).

To determine the market power parameters for Korean milk processors, this paper uses the conjectural variation (CV) concept, with the autoregressive distributed lags (ARDL) approach. This approach fits the data and can be applied regardless of whether the time series are purely I(0), purely I(1) or fractionally integrated. In addition, it allows the estimation of long and short-run market power parameters using the ordinary least squares (OLS). Based on these parameters, the oligopoly power of domestic milk processors for the white fluid milk product is measured.

Following the approach used by Ahn (2006) and Jeon (2009) the model is derived as an application of the ARDL bounds testing. This study focuses on white fluid milk products, which are assumed to be a homogenous product, with *n* firms. The amount of domestic consumption of fluid milk is given by $Q = \sum_{i=1}^{n} q_i$, where q_i is the quantity of milk supplied by processor i. The market demand faced by the domestic milk processors is given by:

$$Q = Q(P_r, Z), \tag{4}$$

where P_r is the retail price of white fluid milk and *Z* is a vector of demand shifting variables.

Given this demand function, the profit function faced by a representative domestic milk processor *i* can be represented as,

$$\pi_i = P_r(Q)q_i - (P_f(Q_f) + \boldsymbol{w})q_i - FC_i, \tag{5}$$

where $P_r(Q)$ is the inverse demand function for white fluid milk, $P_f(Q_f)$ is the inverse demand function for raw milk (material input), q_i is the quantity of fluid milk sold by processor *i*, $\boldsymbol{w}$ is a vector of non-material input prices, and FC_i is the fixed cost for the processor *i* to produce white fluid milk.

The raw milk purchase price $P_f\left(Q_f\right)$ can be expressed as $\bar{P}_g$ because the raw milk price is not determined by the interaction of producers and buyers (dairy processors) in the domestic raw milk market but it is given to the processor by the government. Therefore, the first order condition to maximize dairy processor *i*'s profit can be represented as,

$$\frac{\partial \pi_i}{\partial q_i} = \frac{\partial P_r(Q)}{\partial Q} \frac{\partial Q}{\partial q_i} \frac{q_i}{Q} Q + P_r(Q) - \left(\bar{P}_g + \boldsymbol{w}\right) = 0, \tag{6}$$

where $\lambda_i = \frac{\partial Q}{\partial q_i} \frac{q_i}{Q}$ is a conjectural elasticity which is the degree of market power of the white fluid milk processor *i*. If $\lambda_i = 0$, the industry is perfectly competitive, and if $\lambda_i = 1$, the industry is a monopoly or the processors are perfectly collusive. Thus, the parameter of oligopolistic market power, λ_i, lies between 0 and 1.

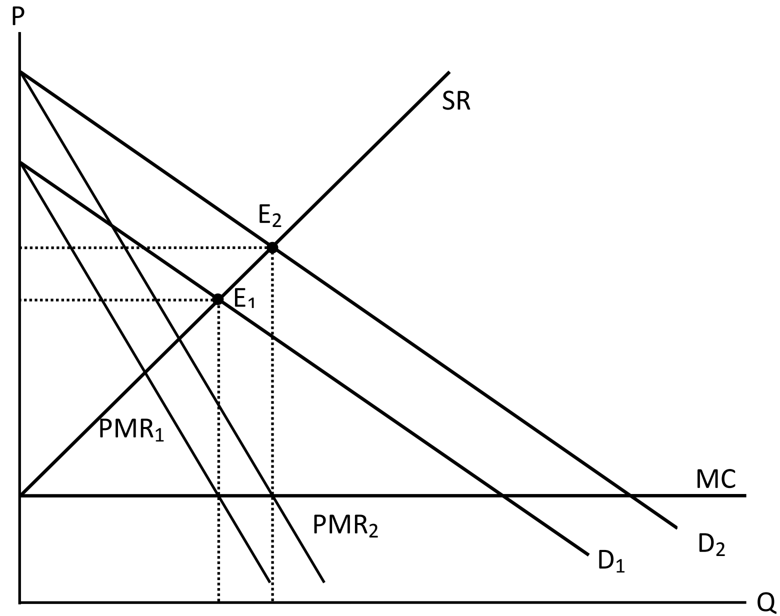
Summing up the first order conditions and then divided by the number of processors *n*, equation (6) can be rewritten as

$$P_r(Q) = -\frac{\partial P_r(Q)}{\partial Q}\theta Q + \left(\bar{P}_g + \sum_{i=1}^{n}\frac{\boldsymbol{w}}{n}\right) \quad (7)$$

where $\theta = \sum_{i=1}^{n}\frac{\lambda_i}{n}$, which can be interpreted as the average market power of white fluid milk processors (Ahn, 2006). Equation (7) includes the slope of the demand curve for white fluid milk, $\frac{\partial P_r(Q)}{\partial Q}$, which shows the relationship between a supply relation and a demand curve. The supply relation may be a supply function, a solution of $P = MC$ but usually the supply relation arises from a market where the sellers may have some market power (Bresnahan, 1982), where it takes the form PMR = MC, perceived marginal revenue equals marginal cost.

Figure 3.1 shows how the supply relation can be derived from the market demand and marginal cost curves. The representative firm in the oligopolistic market faces demand curve D_1, where the demand curve is linear, so the perceived marginal revenue curve PMR_1 is also linear but twice as steep and MC is a constant marginal cost. Typically, the quantity supplied by the firm is decided at the point, where PMR_1 meets MC and the price is decided at the point E_1. This indicates that the quantity and price observed in the market are decided at a point on the demand curve. Nevertheless, the supply curve can be derived with observed price and quantity in the market because the demand curve changes as exogenous variables change. For example, when the demand curve shift from D_1 to D_2 due to the effect on an exogenous variable, the firm decides the quantity at the point, PMR_2 meets MC and the new equilibrium moves to the point E_2. In fact, the line connecting two points, E_1 and E_2, is the supply relation which has a positive slope in price P (Bresnahan (1982) and Ahn (2006)). As we know from the equation (7), the slope of supply curve is expressed as $-\frac{\partial P_r(Q)}{\partial Q}\theta$. In order for it to be valid it has to have a positive sign (+) so there is a negative sign in front because the slope of demand curve ($\frac{\partial P_r(Q)}{\partial Q}$) is also negative. As a result, measuring the degree of market power for white fluid milk processors is identical to estimating the supply relation in white fluid milk market.

Figure 3.1. Static equilibrium representation of the oligopoly market



Source: Bresnahan, Timothy F. (1982) "The Oligopoly Solution Concept is Identified", Fig. 1.

This model assumes that the marginal cost curve is eliminated from every exogenous impact. Therefore, the equation for estimating the supply relation curve should include all possible factors which lead to change in the marginal cost curve. To do this, equation (7) is specified as,

$$P_r = \alpha_0 + \beta_1 Q + \beta_2 \bar{P}_g + \sum_j \alpha_{wj} w_j, \quad (8)$$

where the coefficient β_1 can be expressed as the product of the demand curve slope, $-\frac{\partial P_r(Q)}{\partial Q}$, and the market power parameter θ. This study assumes that the marginal cost of representative firm is constant. Thus, to determine if the domestic milk processors exert their market power in the Korean dairy industry is directly related to β_1 being statistically significant or not.

Among the various factors which affect the marginal cost of white fluid milk, this study considers four major nonmaterial input variables: wage level of manufacturing, packaging cost for white fluid milk, electric charges, and interests rates. The basic empirical model to estimate fluid milk processors' market power can be specified as,

$$P_r = \alpha_0 + \alpha_1 t + \beta_1 Q + \beta_2 \bar{P}_g + \beta_3 wa + \beta_4 pa + \beta_5 el + \beta_6 r + \mathcal{E} \quad (9)$$

where t is a time trend, wa is an average wage of manufacturing, pa is packaging cost, el is electric charges, r is the interest rate, and $\mathcal{E}$ is the random error.

Based on the empirical model, this study estimates the long-run and short-run parameters using an autoregressive distributed lags (ARDL) bounds testing approach and then measures the degree of the white fluid milk processors' oligopoly power.

The ARDL bounds testing approach involves the following two steps (Pesaran et al., 2001). The first step is to examine the existence of a long-run relationship among the variables of interest using an F-test. Once long-run cointegration is established, the second step is to estimate the long-run and short-run coefficients of the same equation.

To examine the existence of long-run relationship, this study uses a more general formula of the conditional ARDL-ECM (Error Correction Model) with an unrestricted intercept and a trend (Pesaran et al., 2001).

$$\begin{aligned}\Delta Pr_t = {} & \alpha_0 + \alpha_1 t + \textstyle\sum_{i=1}^{p} \gamma_{1i} \Delta Pr_{t-i} + \sum_{i=0}^{p} \gamma_{2i} \Delta Q_{t-i} + \sum_{i=0}^{p} \gamma_{3i} \Delta \bar{P} g_{t-i} + \sum_{i=0}^{p} \gamma_{4i} \Delta wa_{t-i} \\ & + \textstyle\sum_{i=0}^{p} \gamma_{5i} \Delta pa_{t-i} + \sum_{i=0}^{p} \gamma_{6i} \Delta el_{t-i} + \sum_{i=0}^{p} \gamma_{7i} \Delta r_{t-i} \\ & + \pi_1 Pr_{t-1} + \pi_2 Q_{t-1} + \pi_3 \bar{P} g_{t-1} + \pi_4 wa_{t-1} + \pi_5 pa_{t-1} + \pi_6 el_{t-1} + \pi_7 r_{t-1} + v \quad (10)\end{aligned}$$

where π_i are the long-run multipliers, γ_i are the short-run dynamic coefficients, Δ is the difference operator and p is the optimal lag length selected by some suitable criteria, and v is white noise error term.

The null hypothesis of no cointegration in the long-run relationship, $H_0: \pi_1 = \pi_2 = \pi_3 = \pi_4 = \pi_5 = \pi_6 = \pi_7 = 0$ in equation (10) is tested against the alternative hypothesis of cointegration, $H_1: \pi_1 \neq 0, \pi_2 \neq 0, \pi_3 \neq 0, \pi_4 \neq 0, \pi_5 \neq 0, \pi_6 \neq 0$ or $\pi_7 \neq 0$ in equation (10) by means of a partial F-test. This F-statistic is a non-standard asymptotic distribution regardless of whether the variables are $I(0)$ or $I(1)$.

The lower and upper bound critical values are used for examining long-run cointegration, where the lower bound critical values assume that all variables are $I(0)$, while the upper bound critical values assume all variables are $I(1)$ (Payne et al., 2011). If the computed F-statistic from equation (10) is higher than the upper bound critical value, the null hypothesis of no cointegration can be rejected, while if the calculated F-statistic is less than the lower bound critical value, the null hypothesis of no cointegration cannot be rejected. If the calculated F-statistic is between the lower and upper bound critical value, the result is inconclusive. In this paper, the critical values reported by Narayan(2005) are utilizing rather than taking those estimated by Pesaran et al.(2001) because our model has relatively smaller sample size, which are shown in Table 2.5.

One of the most important issues in applying the ARDL bounds testing approach is to determine the appropriate lag length p of the first differenced variables in equation (10). Hahmani-Oskooee and Bohl (2000) and Bahmani-Oskooee and Ng (2002) have shown that the results for a long-run relationship among the variables of interest are sensitive to the order of the vector autoregression (VAR). To determine the optimal lag length for the first differenced variables in equation (10) and whether a deterministic linear trend is required, this study used various criteria: Akaike Information Criterion (AIC), Schwarz Bayesian Information Criterion (SBIC), and Hannan-Quinn Information Criterion (HQIC) allowing a maximum lag length of 6. The results are shown in Table 3.1. Under the vector autoregressive model (VAR) the lag order 5 was selected by AIC and HQIC when there is no deterministic linear trend, while with deterministic linear trend the optimal lag order of 1 was selected by all the criteria.

Table 3.1. Lag order selection criteria and asymptotic critical value bounds

Lag order selection under Vector Autoregressive (VAR) model

	Without deterministic trend			With deterministic trend		
Lag	AIC	HQIC	SBIC	AIC	HQIC	SBIC
0	81.98	82.05	84.15	88.15	88.23	88.34
1	68.14	68.68	**69.49***	**6.81***	**7.51***	**8.54***
2	67.50	68.53	70.02	8.76	10.00	11.83
3	57.23	68.73	70.93	9.49	11.28	13.91
4	66.59	68.57	71.47	9.97	12.31	15.73
5	**65.95***	**68.41***	72.00	10.12	13.00	17.22
6	66.01	68.94	73.23	10.72	14.15	19.17

Asymptotic critical value bounds for F-statistic

Prob	0.10		0.05		0.01	
K=6	*I*(0)	*I*(1)	*I*(0)	*I*(1)	*I*(0)	*I*(1)
Case III	2.236	3.381	2.627	3.864	3.457	4.943
Case V	2.657	3.776	3.077	4.284	4.000	5.397

Source: Asymptotic critical value bounds are obtained from Narayan's paper (2005): pp. 1988, 1990

Note: These critical values in Table 2.5 are based on 80 observations. *k* is the number of independent variables per dependent variable in the ARDL models. Case III, and V imply the conditional equilibrium correction models (ECM) with unrestricted intercepts and no trends, and with unrestricted intercepts and unrestricted trends.

* indicates lag order selected by each criterion. AIC, HQIC, and SBIC are Akaike Information Criterion, Hannan-Quinn Information Criterion, and Schwarz Bayesian Information Criterion, respectively.

Chapter 4. Data and Empirical Results

4.1. Description of Data

This study used quarterly industry data from 1985 to 2014. For the key variables, white fluid milk retail prices (P_r) and raw milk prices ($\bar{P}_g$), the Consumer Price Index (CPI) and Producer Price Index (PPI) are used obtained from the Korea Statistical Information Service (KOSIS). These indices were transformed into prices using the prices in 2010 as the basis price[4]. The quarterly consumption of white fluid milk (Q) came from the Korean Dairy Committee (KDC). This consumption is assumed to equal production because of the difficulties of cold storage in South Korea. Wages (wa) are the average salary of workers in manufacturing, which was obtained from the Ministry of Employment and Labor (MOEL). Producer Price Indices (PPI) of paper packaging for food and electric charges are used for data on packaging cost (pa) and electric rates (el), respectively. An interest rate (r) represents the quarterly interest rates. PPI and r were obtained from the Economic Statistics System (ECOS) established by the Bank of Korea.

To remove the effects of general price level changes over time, all price variables used for estimating the market power parameters were deflated by the Consumer Price Index(CPI) of total items (2010=100). Table 4.1 gives the variable name, description, source and descriptive statistics for all of the variables used in the model.

[4] Korean Dairy Committee reported the retail price of white fluid milk and raw milk price paid to the dairy farmers were 2,111 won/liter and 855 won/liter in 2010, respectively.

Table 4.1. Description and descriptive statistics for variables and data source

Variable	Description	Units	Mean	Std. Dev	Source
P_r	Retail price of white fluid milk based on CPI	won/ℓ	1,719.3	271.0	KOSIS
Q	Total consumption of white fluid milk	ton	319,552.7	50,256.7	KDC
$\bar{P}_g$	Raw milk price based on PPI	won/ℓ	867.6	98.5	KOSIS
wa	Real average wage in manufacturing	1,000 won	2,127.1	767.0	MOEL
pa	PPI of paper container for food	Index	101.1	19.4	ECOS
el	PPI of electric charges	Index	112.5	29.7	ECOS
r	Interest rate	%	9.6	4.8	ECOS

Source: KOSIS (Korea Statistical Information Service), KDC (Korean Dairy Committee), MOEL (Ministry of Employment and Labor), and ECOS (Economic Statistics System)

4.2. The Results of the Unit Root Test

The formal unit root tests were performed on all of the endogenous and exogenous variables using the augmented Dickey-Fuller (ADF), Phillips-Perron (PP), and DF-GLS tests. The unit root tests were also performed on the first differenced variables. The results and the critical values for each test are shown in Tables 4.2 and 4.3. If the computed absolute value of the Z-statistics and tau-statistics exceeds each critical value, we can reject the null hypothesis that the time series is nonstationary.

In the ADF test and PP test, only total consumption of white fluid milk (Q) in each test rejects the null hypothesis at the 5% significance level, indicating the time series is stationary, that is, it is *I(0)*. The time series for the average wage (wa) and interest rates (r) reject the null hypothesis in the model with a constant and a time trend; wage and interest rates are rejected in PP test at the 5% and 10% significance level, respectively. Retail price (P_r), raw milk price $(\bar{P}_g)$, and packaging cost (pa), fail to reject the null hypothesis in both cases, implying there is a unit root or nonstationarity, $I(1)$.

In the DF-GLS test with a linear time trend, all the variables cannot reject the null hypothesis at the 5% significance level, meaning these time series are nonstationary, *I(*1*)*. The results from the three different unit root tests suggest that the time series of the variables for estimating market power of Korean milk processors are fractionally mixed, *I(*0*)* and *I(*1*)*.

Table 4.2. ADF, PP, and DF-GLS unit root tests

Variable	ADF test (Z-statistic)		PP test (Z-statistics)		DF-GLS (tau)	Lags
	Cons	Cons & *t*	Cons	Cons & *t*		
5% critical value	-2.889	-3.448	-2.889	-3.447	-2.994	
Retail price (P_r)	-0.229	-1.877	-0.479	-1.945	-1.727	2
Total consumption (Q)	**-4.335****	**-4.596****	**-5.388****	**-6.343****	-1.383	4
Raw milk price ($\bar{P}_g$)	-2.106	-1.372	-1.936	-1.163	-0.875	4
Wage (wa)	-1.135	-2.667	-1.065	**-6.136****	-2.416	4
Packaging cost (pa)	-2.426	-2.109	-2.428	-1.917	-1.442	1
Electric charges (el)	**-3.504****	-2.528	**-4.192****	-1.795	-0.704	3
Interest rate (r)	-0.966	-3.101	-1.171	**-3.159***	-2.306	2

Note: ** and * imply stationarity of the variable based on the specific test at 5% and 10% significance level, respectively.
"Cons" and "Cons & *t"* refer to a model with a constant and with a constant and a time trend, respectively.
The lag lengths of the augmented autoregressions were selected based on the modified AIC of Ng and Perron, (2001).

In general, if a time series has a unit root, the first differences of such time series are stationary (Gujarati, 2003). Table 4.3 shows that the first differences of all the time series are *I(0)* for the ADF and PP tests. However, the DF-GLS test was *I(1)* for the total consumption of white fluid milk. Despite of the nonstationarity of total consumption in the DF-GLS test, the result in Table 4.3 is sufficient to support the fact that the first differences of the times series are stationary.

Table 4.3. ADF, PP, and DF-GLS unit root tests **with first-differenced variables**

First-differenced variable	ADF test (Z-statistic)		PP test (Z-statistics)		DF-GLS (tau)	Lags
	Cons	Cons & *t*	Cons	Cons & *t*		
5% critical value	*-2.889*	*-3.448*	*-2.890*	*-3.449*	*-2.995*	
Retail price (P_r)	**-8.215****	**-8.337****	**-8.574****	**-8.607****	**-8.234****	1
Total consumption (Q)	**-4.268****	**-4.438****	**-22.185****	**-22.865****	-1.906	3
Raw milk price ($\bar{P}_g$)	**-4.805****	**-5.402****	**-10.724****	**-11.216****	**-4.446****	3
Wage (wa)	**-4.768****	**-4.812****	**-22.051****	**-22.117****	**-4.648****	3
Packaging cost (pa)	**-6.736****	**-6.910****	**-8.194****	**-8.307****	**-7.005****	1
Electric charges (el)	**-3.019****	**-4.485****	**-5.750****	**-7.788****	**-4.113****	2
Interest rate (r)	**-8.515****	**-8.506****	**-7.387****	**-7.373****	**-8.464****	1

Note: ** implies stationarity of the variable based on the specific test at 5% significant level and "Cons" and "Cons & t" implies constant and constant with a trend, respectively.
The lag lengths of the augmented autoregressions were selected based on the modified AIC of Ng and Perron, (2001).

4.3. The Results of Cointegration Test: EG and AEG Tests

To test whether two $I(1)$ series are cointegrated, the Engle-Granger (EG) and augmented Engle-Granger (AEG) tests examine the residuals of the DF and ADF unit root tests. The results of these two methods are shown in Table 4.4. The EG test shows that all of the residuals from the regression of white fluid milk retail price (P_r) on each explanatory variable are $I(1)$ at the 5% significance level, which implies that we fail to reject the null of no cointegration between them. These results are also shown in the AEG test but the residual from the regression of the retail price (P_r) on the average wage (wa) is $I(0)$ at the 10% significance level, which suggests they are cointegrated.

Table 4.4. Engle-Granger (EG) and Augmented Engle-Granger (AEG) cointegration tests

		Engle-Granger (Z-statistic)	Augmented Engle-Granger (Z-statistic)	Lags
10% critical value		*-1.611*		
5% critical value		*-1.950*		
1% critical value		*-2.597*		
Dependent	**Retail price ($\boldsymbol{P_r}$)**			
Independent	Total consumption (Q)	-0.936	-0.241	3
	Raw milk price ($\bar{P}_g$)	-0.399	-0.403	3
	Average wage (wa)	-1.581	**-2.019***	2
	Packaging cost (pa)	-0.468	-0.617	3
	Electric charges (el)	-0.126	-0.152	3
	Interest rate (r)	-1.596	-1.442	3

Note: * imply at the 10% significance level.
The lag lengths of the AEG test were selected based on Akaike Information Criterion (AIC).

4.4. The Results of ARDL Bounds Testing

The results obtained from the pre-tests, unit roots and cointegration tests, indicates that not only most variables included in the model, but also their linear combination are nonstationary, which might produce a spurious regression problem. To examine the existence of a long-run relationship among the variables, an ARDL bounds test for cointegration is performed based on the equation (10).

Table 4.5. provides F-statistics for the test of the presence of long-run relationships among the variables and those values depend on the selection of the lag length p. The computed F-statistics in Table 4.5 are compared with the critical value bounds reported in Table 3.1, where the critical values for Case III and V at the 1% significance level for Case III and V are (3.457,

4.943) and (4.000, 5.397), respectively[5]. In the model for p=0 to p=3 with a deterministic trend, the null hypothesis that there exists no long-run relationship or no cointegration is rejected at the 1% significance level, irrespective of whether the regressors are purely $I(0)$, pure $I(1)$ or mutually cointegrated because the calculated F-statistics exceed the upper bound of the critical value. For p=4 the bounds tests is inconclusive. For p=5, the null hypothesis of no level relationship is not rejected since the statistic lies below the 0.01 lower bound. When the bound test is applied to the model without a linear trend, for p=0 to p=3, the null of no cointegration is conclusively rejected, while for p=4 and p=5, the F-test is inconclusive. Therefore, if a linear trend is not included, the bound F-test does not reject the null even if p=4 and 5, which implies that the selection of lag order p=5 based on AIC and HQIC would not be suitable when estimating short-run dynamics. Overall, these test results suggest that there exists a long-run relationship among the variables when a relatively lower lag order, p=0 to 3 was selected rather than a higher lag order, p=4 and 5.

[5] Under the 0.10 and 0.05 critical value bounds we reached the same conclusion in both cases, where for all lag-length the null hypothesis of no cointegration is rejected, while we figured out more useful results under the 0.01 critical value bounds.

Table 4.5. F-statistics of ARDL bounds testing

	Without a deterministic trend	With a deterministic trend
Lag-Length (*p*)	F_{III}	F_V
0	**10.50**[c]	**10.13**[c]
1	**7.68**[c]	**7.35**[c]
2	**7.07**[c]	**6.74**[c]
3	**5.66**[c]	**5.45**[c]
4	4.67[b]	4.50[b]
5	3.89[b]	3.76[a]

Note: In the ARDL bounds testing, the lag length *p* is minus one of the lag length chosen from Table 3.1 because the conditional error correction model in equation (10) uses the first-difference variables, unlike the Vector Autoregressive model (VAR) is used to decide the lag length in Table 3.1.
F_{III}, and F_V imply the F-statistics of the model with unrestricted intercepts and no trends, and with unrestricted intercepts and unrestricted trends, respectively.
[a] indicates that the statistic lies below the 0.01 lower bound, [b] that it falls within the 0.01 bounds, and [c] **that it lies above the 0.01 upper bound.**

4.5. Empirical Results for Estimating Market Power of Korean Milk Processors

After the long-run cointegration relationship was established, the next step was to estimate the long-run and short-run parameters using the ordinary least squares (OLS) technique. Contrary to expectation that including a linear trend may improve the fit of the data, it turned out that the deterministic trend term was not statistically significant, hence, that term was excluded from the empirical regression model. Based on these results, the market power parameters for Korean white fluid milk processors were estimated.

4.5.1. ARDL long-run estimation

The results from the long-run model estimated using the ARDL approach are presented in Table 4.6. It shows that the coefficients of all the explanatory variables were statistically significant at or above the 10% level and the R^2 was 0.85. In addition, most of the estimated coefficients and signs seem to be consistent with the classical economic theory. The coefficient for the total consumption of white fluid milk (Q_t), β_1, was 0.0008 and statistically significant at the 5% level. This implies that the $-\frac{\partial P_r(Q)}{\partial Q}\theta$ term in the equation (7) was positive because the slope of the demand curve was negative. The coefficient for the raw milk price ($\bar{P}g_t$) was 2.23 and statistically significant at the 1% level, meaning an increase in the raw milk price of 1 won/ℓ leads to an increase in the retail price of white fluid milk of more than 2.23 won/ℓ, on average. For the nonmaterial input variables, average wages (wa_t), packaging costs (pa_t), and interest rates (r_t), all of the coefficients have positive signs and were statistically significant. The interpretation of the positive relationships between retail price and the nonmaterial inputs was straightforward because a vertical increase in the marginal cost (MC) curve caused by the changes of input prices results in an increase of the retail price at equilibrium. However, the negative sign of the other input price (el_t) was not expected.

Table 4.6. ARDL long-run estimates

Dependent variable	Retail price of white fluid milk (Pr_t)	
Independent variables	Co-efficient	*t*-statistics
Total consumption of white fluid milk (Q_t)	0.0008**	2.36
Raw milk price ($\bar{P}g_t$)	2.238***	9.53
Average wage (wa_t)	0.364***	8.15
Packaging costs (pa_t)	2.820**	2.01
Electric charges (el_t)	-2.741**	-2.14
Interest rates (r_t)	9.067*	1.75
Constant	-1304.055***	-5.46
Observations	120	
R-square	0.85	

Note: *, ** and *** denote 10%, 5% and 1% level of significance, respectively.

4.5.2. Estimation of short-run dynamics

The final step of ARDL bounds testing approach was to estimate the short-run coefficient. The conditional error correction model (ECM) regression associated with this step is shown in Table 4.7. The lagged error correction term (ect_{t-1}), one-period lagged equilibrium error term from the long-run estimation, and its coefficient was negative and statistically significant at the 5% level, confirming that a long-run equilibrium relationship exists among the variables (Bannerjee et al., 1998). The lagged error correction coefficient was -0.10, which means the deviation from the long-run equilibrium was corrected by 10% by the next period. The short-run changes in total consumption (ΔQ_t) have a statistically positive impact on the short-run changes in the retail price of white fluid milk (ΔPr_t). In addition, the coefficient of the short-run changes in raw milk price ($\Delta \bar{P} g_t$) was positive and statistically significant at the 1% level, indicating a positive relationship between short-run changes in raw milk and retail prices. For the short-run changes of nonmaterial input prices, only the coefficient of average wage (Δwa_t) was statistically significant at the 1% level, implying a positive influence on the short-run changes in the white fluid milk retail price.

Table 4.7. Short-run estimates from the ARDL (0, 0, 0, 0, 0, 0, 0) specification

Dependent variable	Retail price of white fluid milk (ΔPr_t)	
Independent variables	Co-efficient	t-statistics
Total consumption of white fluid milk (ΔQ_t)	0.0002**	2.02
Raw milk price ($\Delta \bar{P}g_t$)	1.605***	9.90
Average wage (Δwa_t)	0.075***	2.71
Packaging costs (Δpa_t)	1.849	1.40
Electric charges (Δel_t)	-0.552	-0.31
Interest rates (Δr_t)	1.958	0.49
Error correction term (ect_{t-1})	-0.103**	-2.52
Intercept	5.514	1.65
Observations	119	
R-square	0.56	

Note: The intercept in the first-difference model represents the coefficient of the trend variable.
** and *** denote 5% and 1% level, respectively.

4.5.3. Oligopoly power of Korean milk processors

The key finding of this study was that the coefficients for white fluid milk consumption in both long-run and short-run dynamics were positive and statistically significant. Those parameters include the market power parameter θ, which can be interpreted as the average market power exerted by domestic white fluid milk processors. However, to obtain the market power parameter, θ, we need additional information regarding the slope of the demand curve $-\frac{\partial \mathrm{P_r(Q)}}{\partial \mathrm{Q}}$. Since the slope of the demand curve is not known, the price elasticity of demand for white fluid milk was assumed to be the average over the entire study period, making the market power parameter θ (Ahn, 2006 and Jeon, 2009):

$$\theta = \beta_1 * (-\frac{\partial \mathrm{Q}}{\partial \mathrm{P_r(Q)}}) = \beta_1 * \eta * (\frac{\bar{Q}}{\bar{P}_r}) \tag{11}$$

where η was the price elasticity of demand for white fluid milk, and $\bar{Q}$ and $\bar{P}r$ were the mean values of consumption and retail price for white fluid milk over the same periods. The domestic white fluid milk market is perfectly competitive if $\theta = 0$, and a monopoly if $\theta = 1$.

As shown in equation (11), calculating the market power parameter using this method is sensitive to the price elasticity of demand; that is, as the price elasticity increases, the market power parameter also increases. This study does not directly estimate the demand function for white fluid milk, but the Korean price elasticity of demand reported by previous research ranged from -1.47 ~ -0.23, where the estimates by author are: -0.57 (Lee, 1997); -0.33 ~ -0.23 (Baeck and Lee, 2002); -1.47 (Shin and Jeong, 2003); -0.96 (Song et al., 2005); and -1.37 (Lee et al., 2005). Given this range, sensitivity analysis was used for various price elasticities of demand and the results are shown in Table 4.8.

In the long-run, when the average price elasticity of demand for the study period was used, the calculated market power parameters were 0.074 for relatively inelastic demand ($\eta = -0.5$), 0.149 for unit elastic demand ($\eta = -1.0$), and 0.223 for relatively elastic demand ($\eta = -1.5$). In comparison to the previous studies (Ahn, 2006 and Jeon, 2009), the estimates for oligopoly power were similar. The computed short-run market power parameters were 0.019 for relatively inelastic demand ($\eta = -0.5$), 0.037 for unit elastic demand ($\eta = -1.0$), and 0.056 for relatively elastic demand ($\eta = -1.5$).

Table 4.8. Oligopoly power under various price elasticities of demand for white fluid milk

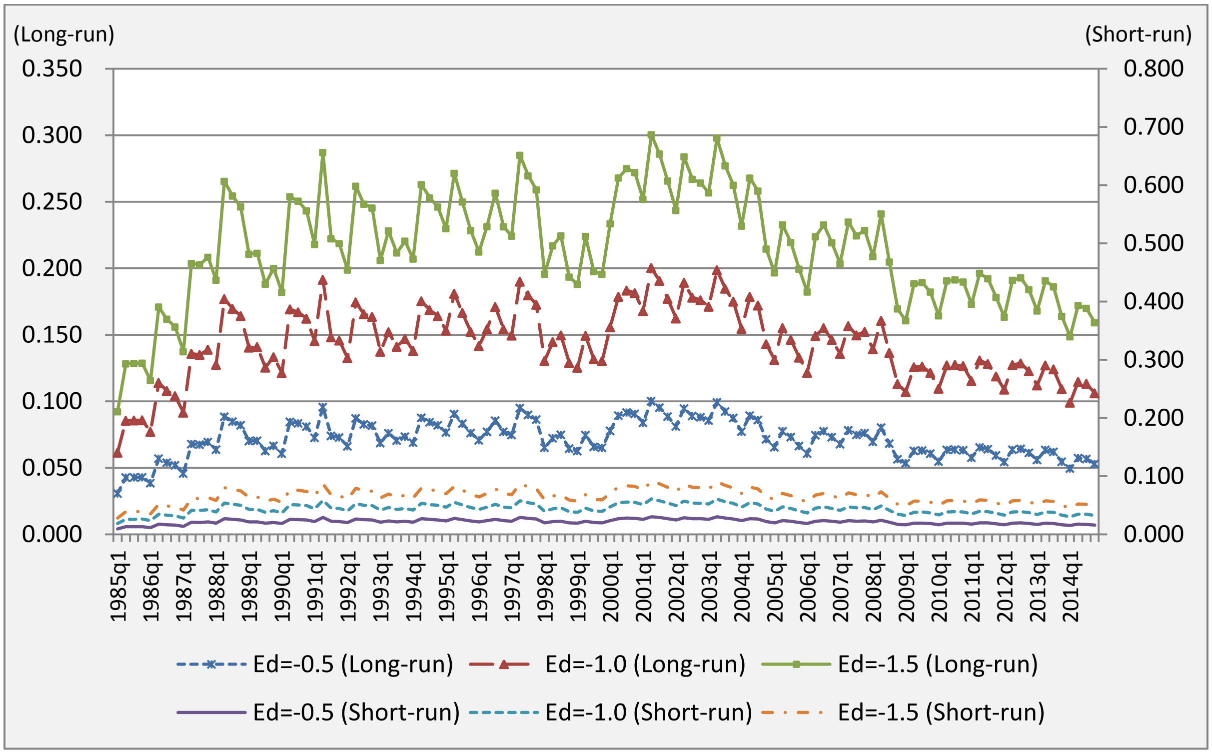
		White fluid milk					
		Long-Run			**Short-Run**		
$\hat{\beta}_1$		0.0008			0.0002		
η		Inelastic	Unit elastic	Elastic	Inelastic	Unit elastic	Elastic
		-0.5	-1.0	-1.5	-0.5	-1.0	-1.5
θ		**0.074**	**0.149**	**0.223**	**0.019**	**0.037**	**0.056**
	Ahn (2006)	0.077	0.154	0.232	-	-	-
	Jeon (2009)	0.057	0.115	0.173	-	-	-

Note: The average consumption ($\overline{Q}$) and retail price ($\overline{P}_r$) for white fluid milk are 319,552.7 ton and 1,719.3 won/ℓ, respectively.

Figure 4.1 shows historical changes of the market power parameters in the long and short-run under various price elasticities assuming that these price elasticities were constant throughout the study period. In both the long and short-run, all the parameters have an upward trend until 2003, with steeper increasing ratios from 1985 to 1988, and then a downward trend afterward[6]. This indicates that the white fluid milk processors exerted more oligopoly power in the Korean milk industry until 2003 but then steadily exerted less market power afterward. In terms of changes of market power parameters over time, this result differs from Ahn (2006), which reported that these oligopoly power parameters consistently increased until 1988 but then they decreased until 2004. This previous study also claimed that this market power parameter changing point was substantially coincidental with that of the structural change in the Korean fluid milk consumption[7].

[6] Korea suffered a foreign-exchange crisis in 1997, which resulted in a sharp decrease in the white fluid milk consumption (Song et al., 2005); hence the transitory drop of oligopoly power in the late 1990s would be caused by the financial crisis.
[7] Baeck et al. (2002) and Jeong et al. (2003) determined that the continuously increasing trend in fluid milk demand met a turning point in 1988 and 1989, respectively but then it stagnated.

Figure 4.1. Short and long-run market power parameters under the various price elasticities of demand



To test whether there was a structural break in the relationship between the dependent variable and the independent variables, the sample data were divided into two time periods in the long-run and short-run models: 1985q1 to 2003q4 and 2004q1 to 2014q4; the pre-and post-2003 turning point. The idea behind the Chow test is that if there is no structural change between the two sets of sample periods, then the restricted residual sum of squares (RSS_R) and the unrestricted residual sum of squares (RSS_UR) should not be statistically different. The Chow test statistic is

$$F = \frac{(\text{RSS}_\text{R} - \text{RSS}_\text{UR})/k}{(\text{RSS}_\text{UR})/(n_1 + n_2 - 2k)} \sim F_{[k,(n_1+n_2-2k)]}$$

where RSS_R is the restricted residual sum of squares from the pooled regression, RSS_UR is the sum of squared residuals from each sample period, n_1 is the number of observations in 1985q1 to

2003q4, n_2 is the number of observations in 2004q1 to 2014q4, and k is the number of parameters estimated. The F critical value is from the F distribution with k and $(n_1 + n_2 - 2k)$ degrees of freedom in the numerator and denominator, respectively. If the Chow test statistic exceeds the critical F value, the H_0 for no structural change is rejected.

Table 4.9 shows the calculated F values in the long-run and short-run. In the long-run model, computed F statistic was 19.93, which implies that the H_0 of no structural change can be rejected. This implies that the values of the parameters of the model do not remain the same through the entire time period. The Chow test supports the findings that the oligopoly power parameters have a different trend before and after 2003. However, in the short-run dynamic model the $H_0$was not rejected because the test statistic was less than the critical F value.

Table 4.9. The Chow-test results of the structural change

	Long-run	Short-run
Computed F value	19.93	0.86
H_0: No structural change	Reject ($> F_{(7,120)} = 2.09$)	Not Rejected ($< F_{(8,120)} = 2.02$)

Note: The critical F values are obtained from *F*-table at 5% level of significance.

This decreasing oligopoly power result is still controversial since the Korean milk industry maintained a high level of market concentration, about 75%-80% of CR4, over the same time. There are several remarks on these counter-intuitive findings. First, the assumption of the constant conjectural elasticity parameter is a limitation which may affect the empirical analysis (Gohin and Guyomard, 2000). The Chow-test result also supports this limitation, providing evidence that the conjectural elasticity parameter changes over time. Secondly, the traditional assumption of the positive relationship between market power and the concentration index is less obvious (Gohin and Guyomard, 2000). In addition, Balra (2000) found that this relationship is more complicated rather than simply a positive linkage, showing the U-shaped relationship between market power and farm size inequality (FSI). Finally, there were several studies that found a declining market power in the food industry while the market concentration ratio increased (Schroeter and Azzam, 1991 and Koontz et al., 1993)[8]. Hamilton and Sunding (1997) examined the marginal effect of a farm supply shift on the market structure of the food processing sector and found that when the farm supply curve shifts out, increasing concentration and decreasing market power are likely to occur simultaneously.

In calculating the Korean milk processors' oligopoly power from equation (11), it was directly influenced by the ratio $(\frac{\bar{Q}}{\bar{P}_r})$, which was average consumption of white fluid milk over the average retail price because we assumed that the price elasticity of demand was constant over time. This decreasing trend is due to the fact that the total consumption of the white fluid milk (Q) stagnated after 2003, while the deflated retail price (P_r) increased during the same period. This stagnation of the white fluid milk consumption was closely linked to the growth of substitute industries. Especially, the fermented milk market which was steadily developing since the 1980s and was the second largest industry in Korea dairy market (Lim, 2007). This stagnant consumption of the white fluid milk might have been caused by a sharp increase in fermented milk consumption as consumers demand for healthy food rose (Baeck et al. 2002). In addition, the deflated retail price of the fermented milk products has consistently decreased by 24.6% over the whole study period, which might encourage consumers to switch from the white fluid milk to

[8] Schroeter and Azzam (1991) proposed a plausible hypothesis that a packer operating a large plant in the U.S. hog-packing industry can obtain significant cost economies by assuring the necessary flow of inputs to operate the plant near capacity and this cost savings may have overwhelmed the incentives for oligopolistic output restrictions. Koontz et al. (1993) stated that increased excess capacity may have increased competition in the fed cattle market since firms attempted to spread fixed costs. Also, increased contractual arrangements with boxed beef purchasers may lead meatpackers to break tacit agreements more often.

fermented milk[9]. In addition to the substitution effect, the white fluid milk consumption is also affected by changes in income. According to the Korean Statistical Information Service (KOSIS), monthly real income per household in Korea has continuously increased since 1990, showing 2.83% annual increase on average. However, a few studies have shown that the income effect on the fermented milk demand was greater than that on the (white) fluid milk demand[10], which implies that the households with an increase in income are more likely to purchase the fermented milk rather than white fluid milk. For these reasons, the white fluid milk processors could not exert the same level of oligopoly power as they did at the beginning of the study period because it became easier for the consumers to switch to substitutes.

Finally, the level at which white fluid milk processors exert their market power on the market was measured over different time periods. To do this, the Lerner's Index was calculated based on the results from Table 4.8, which describes a firm's level of market power by relating price to marginal cost, which can be expressed as:

$$L \text{ (Lerner's Index)} = \frac{P - mc}{P} = -\frac{\theta}{\eta} = -\frac{\partial P_r(Q)}{\partial Q}\frac{Q}{P_r}\theta = \hat{\beta}_1 \frac{\bar{Q}}{\bar{P}_r} \tag{12}$$

where P is the market price set by the firm, mc is the firm's marginal cost, θ is the market power parameter, η is the price elasticity of demand for white fluid milk, $\hat{\beta}_1$ is the estimated total consumption of white fluid milk coefficient, and $\overline{\mathrm{Q}}$ and $\overline{\mathrm{P}}_{\mathrm{r}}$ are the mean values of consumption and retail price for white fluid milk over the same periods, respectively.

The calculated Lerner's Indices are the same as the average computed oligopoly power parameters under the price elasticity of 1 so they have a similar shape, which was shown in

[9] Baeck et al. (2002) reported that the fermented milk and the white fluid milk have a high substitutability and the cross-price elasticity of the white fluid milk demand is 1.48-1.65.
[10] There were several studies that reported the income elasticity of demand for (white) fluid milk to be relatively inelastic (Lee, 1997; Song and Sumner, 1999; Baeck, 2002), while the income elasticity of demand for the fermented milk product was 1.25 (Lee, 1997).

Figure 4.1. These Lerner's Indices, represented by the markup ratio ($P - mc$) to the market price, also measure the margins for the white fluid milk processors. This supports the notion that white fluid milk processors do exert a significant degree of oligopoly power by setting the market price of white fluid milk above the competitive price, as shown in Table 4.10. The average Lerner's Indices during the period from 2001 to 2005 represent the highest value of 17.79% for the long-run and 4.45% for the short-run, but afterward, the Lerner's Indices steadily decreased, indicating the domestic white fluid milk market has become more competitive over time.

Table 4.10. The Lerner's Index of white fluid milk processors in different time periods

Time Period	Lerner's Index	
	Long-run (%)	Short-run (%)
1985-1990	13.23	3.31
1991-1995	16.52	4.13
1996-2000	16.33	4.08
2001-2005	17.79	4.45
2006-2010	14.10	3.52
2011-2014	12.40	3.10
1985-2014	14.90	3.70

Chapter 5. Conclusion

5.1. Summary of the Key Results

This study empirically measured the degree of oligopoly power of domestic white fluid milk processors by using the NEIO (New Empirical Industrial Organization) approach. Generally, even though empirical studies on measuring market power have used time series data such as retail prices of a specific product, demand and supply, and other input prices, they overlooked spurious regression problems caused by using nonstationary time series data. This paper first checks for stationarity using three different unit root tests and cointegration analysis to check for spurious regressions.

The results from the unit root tests showed that the time series variables for estimating Korean milk processors market power were fractionally mixed, implying stationary $I(0)$ and nonstationary $I(1)$ time series data were included in the model. However, although the two time series data were individually $I(1)$, their linear combination can be stationary since the stochastic trends between the two series can be cancelled out. The Engle-Granger (EG) and augmented Engle-Granger (AEG) tests results also indicated that the linear combinations between variables were not stationary, implying that empirical regressions using these variables may produce spurious results.

Based on the unit root and cointegration tests, this study adopted the autoregressive distributed lag (ARDL) bounds testing approach. One advantage of the ARDL approach is that it can be applied irrespective of whether the time series data were purely $I(0)$, purely $I(1)$ or fractionally integrated (Pesaran and Pesaran, 1999; and Bahmani-Oskooee and Ng, 2002). A long-run equilibrium among the variables was found when relatively lower lag orders, p=0 to 3, were selected rather than higher lag orders, p=4 and 5. In the long-run estimation, the coefficients of most explanatory variables were statistically significant at or above the 10% level and most of these coefficients and signs were consistent with economic theory. In the short-run dynamics, the short-run changes in independent variables, total consumption, raw milk price, and average wage have a statistically positive impact on short-run changes in the retail price of white fluid milk.

In both the long and short-run, the parameters have an upward trend until 2003 but afterward have a downward trend until 2014, indicating white fluid milk processors exerted more oligopoly power until 2003 but then less market power steadily afterward. Finally, the measured average Lerner's Indices from 1985 to 2014 were 14.9% and 3.7% in long and short run model, respectively. This implies that the domestic milk processors had on average gained excess profits by exerting their oligopoly power. Like the market power parameters trend, Lerner's Indices are steadily decreasing, indicating the domestic white fluid milk market has become more competitive over time.

5.2. Implications

In economics, a concentration ratio (CR) measures the total market share of a given number of firms in an industry, giving intuitive insight into how concentrated an industry is and therefore it usually is used to show the extent of market control of the largest firms and to illustrate the degree of market power. That is, a high concentration ratio may indicate that a significant amount of market power in the industry. However, high concentration ratios do not necessarily imply market power because there is no direct link between the concentration ratios and the extent of non-competitive behavior. The Korean milk processing case is a primary example of this. That is, even though the Korean milk processing industry had maintained almost 80% level of CR4 (Four-firm concentration ratio) since 2003, the empirical results in this study indicate that the oligopolistic power of white fluid milk processors had steadily decreased during the same time period, implying it was becoming a more competitive market. This finding seems to be more reasonable considering the characteristics of the white fluid milk products, that is, it is more difficult for existing firms to create barriers to entry through the technical development, product differentiation and favorable access to essential raw materials.

On the other hand, from the perspective of the government or a policy maker this high concentration ratio in the domestic dairy industry implies that it might be always possible for a few firms to merge, control their production, or engage in price fixing. Therefore, under the current milk pricing system, the policy makers need to determine if a retail price changes due to the raw milk price changes decided by the government or due to milk processors' markup actions and why.

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