Three Essays on Agricultural Prices, Markets, and Trade

by

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Abstract

This dissertation includes three essays to address price analysis, market response as well as trade effects of the agricultural commodities, in both China and U.S. The first chapter examines the demand impacts following a food contamination with focus on the case of *Sanlu* melamine contamination of dairy products in 2008, China. Evidence is presented on media coverage effects, seasonal factors, time trends, and contemporaneous own- and cross-price effects with respect to dairy food consumptions. The average demand response of powdered milk to media reports was small, especially in comparison to price effects, and to previous estimates of related issues. This average small impact on dairy demand can be attributed to the average amount of adverse news information concerning dairy safety being small as well as short lasting, and furthermore, the media coverage had no significant lagged effect on demand.

The second chapter analyzes the farm-retail price transmission for U.S. whole milk according to the conventional Houck approach and to the von Cramon Taubadel and Loy error correction model (ECM) approach by using monthly data over the period from January 1996 to December 2011. Also, to accommodate to Gardner's (1975) model including demand and supply shifters, the study is examining for marketing margin model and price transmission model. Though the tests agrees to the previous examples that increases in the farm price of milk are passed through to the retail level more fully than were decreases in the farm price of milk, there

is no clear-cut conclusions that can be drawn regarding the effects of market power on the degree of price transmission for U.S. whole milk.

The third chapter focuses on the U.S. agricultural trade against the remaining of the world. The dynamic ARDL model of error correction version is applied, not only investigating if there is J-curve effect in the short-run or not, but also taking a deep analysis for U.S. recession effects and exchange rate as well as income growth effects in the long run on the U.S. trade balance of agricultural commodities which mainly consists of bulk products and high-value products. Our results indicate that there is no significant J-curve effect for three cases, while the long-run effect demonstrates that the domestic currency devaluation is positively related with U.S. agricultural trade balance for bulk, high-value and combined agricultural products, though the high-value products appear the more modest effects compared to the other two. In sum, the real trade-weighted exchange rate is found to be the key determinant of U.S. agricultural trade balance in the long-term, rather than domestic or foreign income. We find that the three categories of agricultural products do indeed respond differently to exchange rate and income. For bulk and high-value products, U.S. exports are highly sensitive to exchange rate and foreign income, while U.S. imports barely respond. For combined agricultural products, on the other hand, U.S. exports respond greatly to exchange rate, and U.S. imports behave significantly with respect to both of changes in exchange rate and foreign income; besides, the 1980s recession had significant effects on U.S. trade balance while the most recent recession had great impact on U.S. imports, showing the U.S. trade with ROW partners was mainly influenced by the two times economic crisis during our sample period.

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Chapter I. Does melamine incident impact Chinese dairy demand?

1.1 Introduction

"False views, if supported by some evidence, do little harm, for everyone takes a salutary pleasure in proving this falseness: and when this is done, one path towards errors is closed and the road to truth is often at the same time opened."

---Darwin, Charles. 1859

Dairy safety concerns have arisen in China during the recent years since the outbreak of contaminated milk products. The melamine crisis occurred in 2008, initially coming mainly from the upstream supplement chain of Sanlu dairy company of China, but later found in a large number of other domestic dairy brands. The World Health Organization referred to the melamine incident as one of the largest food safety events it had to deal with in recent years, and the safety issues of the Chinese food market have received much attention.

Dairy contamination can be a big problem as China has the largest population in the world and the potential negative effect of the melamine incident on consumers' demand involves the dairy group products including fluid milk and dry powdered milk as well as yogurt. The objective of this article is to study whether publicized dairy safety concerns have impacted milk consumption and a model is applied to the incident of melamine contamination of milk in mainland China. A media coverage index was constructed to test for hypotheses about the demand impacts on different dairy products. These hypotheses have significant implications for government and industry strategies for responding to contamination incidents.

The chapter is organized as follows. Section 2 summarizes the historical background of the Chinese dairy market in terms of the developments and challenges. In Section 3, the literature including both food safety incidents and consumption demand systems is reviewed. Section 4 provides a theoretical framework for a milk demand system, which is examined with monthly data of weakly separate dairy products in urban China in Section 5. The empirical results including effects of price and certain demand shifters are interpreted in Section 6 and the last section concludes.

1.2 Historical Background

Dairy products are the products made and processed from fresh cow (sheep) milk as the raw material, including liquid milk, powdered milk, butter, cheese, etc. In urban China, the major dairy products are liquid milk (pasteurized milk, sterilized milk), powdered milk and yogurt. The market scales of other dairy products in China such as cheese and cream are all quite small. Since 1998, China's dairy industry has entered a phase of rapid growth and the annual growth rate had been continuing over 10% for one decade. In 2006, the total dairy consumption reached 19 mmt (million metric tons). Demand from China is already having a bearing on market prices for many dairy goods across the world. In 2008, the per capita consumption of liquid milk arrived at 9.5 kg at the national average level, which was no more than half of the liquid milk consumption, 20.25 kg person for urban China in the same year (Figure 1); however, the demand is still low compared to per capita consumption of other countries such as Japan (34.2 kg), South Korea (35 kg), India (38.2 kg) and the United States (82.3 kg) (FAO, 2009).

Nevertheless, the melamine incident of milk contamination, disclosed in September 2008, dealt a serious blow to China's dairy industry, especially in the end of that year. The level of melamine found in the contaminated infant formula had been as high as 2560 milligrams per kilogram ready-to-eat product, while the level of cyanuric acid was unknown. The chemical was added by the milk producers because it tended to make the dairy products appear to have higher protein content and earn more profits. Initially the sixteen infants who had been mainly fed on powdered milk produced by the Shijiazhuang-based Sanlu Group (partly owned by New Zealand's Fonterra group) were diagnosed with kidney stones. Until November 2008, the media reported estimated 300,000 victims, with six infants dying from kidney stones and other kidney damage. Under pressure of a number of complaints, the Sanlu Group decided to recall all milk products produced prior to Aug. 2008; this company declared bankruptcy in Dec. 2008. Later in Jan. 2009, four major defendants, including a top dairy company executive, were sentenced to life in prison as well as deprived of political rights for life under the draft law; meanwhile, they were imposed a total fine of RMB (Renminbi) 24.7 million yuan.

As seen in Figure 1.1, the average monthly per capita consumption decreased 11.6%, 7.1% and 38.8% for fresh milk, powdered milk and yogurt, respectively from September to October of 2008 after the media coverage exposed the melamine crisis to the public; the correlation between dairy consumption and media coverage appeared very strong within that one month, indicating that news coverage on the melamine incident did impact and undermine the consumers' confidence of purchasing dairy products.

1.3 Literature Review

Previous related studies can be sorted into two categories, the methodology of the dairy demand system and the effect of unfavorable information on the public.

For the demand part of the analysis, Adelaja (1991) presented estimates of a complete demand system for food at home with special emphasis on dairy products with an AIDS (Almost Ideal Demand System) Model and came to the conclusion that the demand for dairy products is generally inelastic, cross-price effects are moderate and income effects are small and negative.

Zhang and Wang (2003) estimated demand elasticities and the impacts of regional and demographic variables for 17 food products through a two-stage budgeting procedure with complete demand systems using datasets of 3,500 urban Chinese households, concluding that education level affects fresh milk consumption negatively and interestingly, household size tends to affect per capita fresh milk consumption positively. They contributed the result to the availability and accessibility to these dairy commodities that are highly associated with the development of infrastructure and channels of production and distribution.

Also, Yen et al. (2004) investigated household food consumption in urban China by using data from the 2000 Survey of Urban Households and concluded that dairy consumption has grown rapidly in China over the last decade, which had increased from 1.37 mmt in 1980 to 12.06 mmt in 2002, with the annual rate of growth reaching about 10.4% for fluid milk consumption (USDA-FAS, 2003). A translog demand system was estimated taking account of reported zero consumption, and derived high expenditure elasticities for milk, suggesting that demand for these products would grow faster than demand for other products as the economy develops and incomes increase. Meanwhile, they demonstrated that demand is more price-responsive for milk than all other food products, and net substitution is observed among most food products.

Meanwhile, Fuller et al. (2007) studied the consumption of dairy products in urban China including Shanghai, Beijing and Guangzhou using micro-survey data. The results from estimation of a double-hurdle model of consumption showed that income and marketing channels are the key determinants of milk consumption levels, and they also concluded that education, advertising, and convenience play important roles in consumption of other dairy products. They figured out that the growing sophistication of China's retail sector is influencing consumption of dairy products.

Recently, Zheng and Henneberry (2009) stated that dairy demand in China is anticipated to increase in the future with growing urbanization and westernization associated with the growth in per capita incomes (Fuller et al., 2007; Yen et al., 2004). The dairy category is more responsive to changes in own price and expenditure than most food products examined in this study. The high expenditure and own-price elasticities (in absolute value) for dairy products show that both income and own price play important roles in dairy food consumption.

However, matters would become more complicated when taking information effects into account. Previous authors have followed several alternative strategies to assess the impacts of information on food safety. The most classical case study was heptachlor contamination of fresh fluid milk in Oahu, Hawaii in 1982 reported by Smith et al. (1988); they followed a second-order Almon polynomial structure that constrained their milk media index to examine the impact of the heptachlor contamination of milk and also estimated the impact in the Hawaiian issue by including a variable related to media coverage in a demand function; the applicable procedure for estimating sales loss following the incident implied that the demand curve for the product would likely shift downward after the incident. Empirical studies have attempted to capture this effect by incorporating dummy and/or trend variables to represent the health warning in models of consumer purchases (Hamilton, 1972; Mowen, 1980; Mowen and Pollman, 1982; Shulstad and Stoevener, 1978; Schuker et al., 1983).

By contrast, Foster and Just (1989) discarded the media variable and substituted for it a nonlinear shift on the intercept that allows for an exponential decrease in the food scare effects and also some long-term persistence. Mazzocchi et al. (2006) measured the time pattern of multiple and resurgent food scares and their direct and cross-product impacts on consumer response. They studied Italian aggregate household data on meat demand and assessed the time-varying impact of a resurgent Bovine Spongiform Encephalopathy (BSE) crisis (1996 and 2000) and the 1999 Dioxin crisis; and the empirical results showed little relevance of the Dioxin crisis in light of the preference shift, while not excluding the more relevant price effect.

Chang and Kinnucan (1991) examined the impact of cholesterol information and advertising, representing the positive and negative information respectively, on consumption of fats and oils (while focusing on butter) in Canada. They concluded that consumers' response to negative information appeared to outweigh their responses to positive information and particularly, the increased consumer awareness concerning health effects might have contributed to the secular decline in butter consumption.

Kinnucan et al. (1997) did research on the effects of health information on U.S. meat demand and suggested that the contemporaneous effects of health information in general are larger than the lagged effects, indicating a relatively rapid decay in health information effects, coupled with the insignificance of many of the lagged terms. Furthermore, their study showed the health information elasticities in general are larger in absolute value than price elasticities, showing that small percentage changes in health information have larger impacts on meat consumption than equivalently small percentage changes in relative prices.

Piggott and Marsh (2004) developed the expanded GAIDS (Generalized AIDS) model accounting for a media index specifically built for the contamination impact on meat demand about Bovine Spongiform Encephalopathy (BSE); they constructed food safety indices separately for beef, pork and poultry allowing for the investigation of separate own- and cross-commodity impacts from food safety concerns and lead to the conclusion that the impact of food safety information on demand was determined to be limited to a contemporaneous effect.

1.4 The Model

Swartz and Strand (1981) developed a model of how consumers respond to a contamination incident in the case of kepone in 1981 and argued that consumer's utility function can be expressed as $U\ (X_i\ (Z_i\ (N)))$.[1] In their study of demand for oysters in Baltimore they found that the measure of negative media coverage of the ban was statistically significant in explaining declines of purchases of oysters. Later in 1982, Smith et al. also applied such a model in the case of heptachlor contamination of fresh milk in Oahu and their focus was to find how a consumer allocated income among goods when information about the quality of one of those goods has changed; in the end, they found that media coverage following the incident had a significant impact on milk purchases.

The dairy category of this article represents a weakly separable food grouping[2] and variable of income should be replaced by total expenditure (TEXP) assigned for the group. P_1, P_2, and P_3, denote respective prices for each product within the group. The demand for X_i is a function of expenditure, prices, and information:

(1) $X_i = X_i\ (P_1, P_2, P_3, \text{TEXP}, N)$

As Smith et al. (1982) noted, information can be classified into favorable versus unfavorable, or positive versus negative, and the unfavorable information may have a greater impact because it is less common in an individual's social environment than the positive cues of favorable information (Kanouse and Hanson, 1971). However, Kahneman (2011) argued, by citing a number of experimental examples, that the boundary between good and bad is a reference point that changes over time and depends on the immediate circumstances. That means

[1] It stands for consumer's perception of the quality (Z_i) of a good (X_i) affects that consumer's level of utility and the perception of product quality will depend on information index (N).

[2] The utility function mentioned above can be partitioned into separate subsets or branches, such as food or, at a more disaggregated level, dairy products; which means that consumers decide first on their total consumption of food, then on the budget allocation among food groups, and finally on the allocation among individual commodities within a specific food grouping. (Pollak, 1971).

at any given moment present condition and past reference both determine the utility function. Consequently, the unfavorable information provided either by the factual performance or vested interest (Mizerski, 1982) would possibly turn into less unfavorable taking the reference point into account, or rather, may be getting worse given the situation that loss-aversion outweighs other positive elements. This two-sided distribution cannot be identified exactly and if ignoring the reference point, the measurement of difference between favorable and unfavorable information would be not accurate.

A couple of related hypotheses accordingly suggest that the melamine incident can be considered as a bad event in the study and therefore, media coverage on such an incident would be the bad (unfavorable) information that was absorbed by consumers and may influence their purchasing behavior as well as demand for dairy food. The extent of dairy demand shift should depend on whether unfavorable information concerning the melamine incident was available to consumers, how likely they received that bad information, and to what degree would their subsequent evaluation be about the tradeoff between loss and net gains of consumption. Further, the Internet news reports are assumed to be the only source of information about melamine contamination based on the availability of data selection, although milk quality can be communicated to consumers from numerous other sources including in-store information and information sent from directly to buyers and the word of mouth (Smith et al., 1988).

Media Coverage Index

Following earlier case of the Hawaii milk contamination incident (Smith et al., 1988), media coverage indices were constructed based on newspaper articles from the popular press. Different from the above study, the electronic news or articles related to the incident were searched by the author from portal websites of China as the Internet-browsing technology has been widely popular in urban cities, rather than the traditional habit of newspaper reading. Thus, data for the series were obtained by searching in each news page of the three major portal websites of *Sina*, *Sohu* and *Fenghuang*, which rank top three with regard to the number of visits over the sample period in all of China. Keywords searched were narrowed to *Sanlu milk melamine incident* or *Sanlu milk contamination*, because this can make selections constrained to barely favorable or the unfavorable information from the websites. The media coverage index was first measured by scaling probability of influencing consumer judgments of milk safety and then weighted by attention score from each news page concerned. Specifically speaking, the media index was constructed by coding articles from the three web portals during the post-incident period, since each website had released a certain number of articles or news each month, and the average scores of 0, .25, .75 or 1 were assigned to articles/news of each website in each month depending on the author's judgment of probability [3] that the reported articles would negatively affect consumers' confidence of consuming milk (Swartz and Strand); then these codes were assigned a scaling weighted number of 0 to 1 by the prominence of monthly website articles using the attention score.[4] The weighted codes collected from the three biggest Chinese websites were finally summed for each month for media coverage index. The larger the number is, the more negative the effect would be. The effective single index of unfavorable information variable for all the three dairy products was then built up as a running total applying the expression

(2) $X_{i,\,t} = \sum \rho_{j,\,t}\, \omega_{j,\,t}$

[3] The judgment was based on the website visits and appearance frequency of key words. For example, if the *Sanlu* Company or government issued public statements such as the problem eliminated or all of the contaminated products recalled, the assigned score would be zero.

[4] This method was imitated by Budd (1964) and the weighted scale ranged from 0 to 5 for newspaper articles, but in this paper, the weighted number is scaling by 0, 0.2, 0.4, 0.6, 0.8 and 1 for electronic news which differentiates from traditional ones. For example, the article appearing in news section of *Sina* website in late year 2009 would be assigned a weight of 0.6.

where $X_{i,\,t}$ is the negative information datum collected from each website (i = 1, 2, 3) in each month (t = 36, . . . ,75), $\rho_{j,\,t}$ is the scaling scores of influencing consumer judgments ranging from 0 to 1 and $\omega_{j,\,t}$ is the weighted attention score ranging from 0 to 1, (j is representing each selected reported article of each website)

(3) $N_t=\sum_{t=1}^{t} X_{1,\,t} + \sum_{t=1}^{t} X_{2,\,t} + \sum_{t=1}^{t} X_{3,\,t}$

where N_t is the sum of negative information data collected from three websites in each month.

Overall, the media coverage of the melamine incident reveals higher concerns for powdered milk than liquid milk or yogurt over the sample period. Figure 1.1 plots the media coverage indices. During the period of *ex ante* incident the news indices remained zero; from Sept. 2008, the media series increased significantly but was starting to wane by the end of 2009, when the indices reached a new, historic high in February and August of 2010 respectively in that the dairy safety issue became the national focus of attention again, pertaining to some similar important events.[5] The sales of dairy products were all greatly affected in Sept. 2008 but only powdered milk was highly related with media index change for each time, showing that as the adverse media reports increased (indices increased), the per capita consumption of powdered milk went down immediately.

Empirical Specification

For the weakly separable grouping of dairy food, the conditional demand equations (Edgerton, 1997; Pollak, 1971) are specified for liquid milk, powdered milk and yogurt; thus following Chang and Kinnucan (1991), a semilogarithmic form was selected in that the utility

[5] In February and August 2010, respective media reported that the *Shengyuan* powered milk contained contents that could cause precocious puberty and the Mengniu milk added some subsistence which might invoke cancer problems. These articles mentioned *Sanlu* incident again, contributing to the increased media indices in those two months.

maximization can be satisfied (Hanemann, 1982) and in accordance with Smith, et al. (1988) the *Almon* polynomial lag model is estimated without log form[6]. The demand model for the three dairy products (fresh milk, powdered milk and yogurt) is set up here[7]. Thus, the estimated equation is

$$(4)\quad Q_{i,t} = \alpha_i + \alpha_j \ln P_{j,t} + \beta_i \ln (\mathrm{TEXP}_t/P_t^*) + \theta_i DV + A(L)N_t + \phi_k S_k + \varepsilon_{i,t}\ ,$$

where t stands for monthly observation and t = 1, . . . ,75 from October 2005 to December 2011; Q_i is monthly per capita consumption of dairy product i (i = 1, 2, 3); P_j is the real price of good j[8]; TEXP is the consumers' total group expenditure on dairy deflated by the Stone price index, $P_t^* = \ w_{j,t} \ln P_{jt}$ (Deaton and Muellbauer, 1908, p. 62), where w_j is the consumer's expenditure share of good j of the dairy products; DV is the dummy variable that is zero before the September 2008 contamination and 1.0 thereafter; N_t measures the negative media coverage which was defined earlier and A(L) is a polynomial lag structure of the media variable; S_k's are seasonal dummy variables (month of December is the omitted category); and ε_i's are random error terms.

1.5 Data

Dairy data used in this article are monthly consumption over the period from 2005 to 2011,[9] providing a total of 75 observations.[10] The basic quantity data are per capita consumption data from the Dairy Association of China (DAC) published in both the China Dairy Yearbook and updated publications available online. The real price of liquid milk is the average monthly retail fresh milk price; and the prices for both powdered milk and yogurt were calculated by their

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[6] The level variable of media indices is used here because there are too many zeros during *ex-ante* incident.

[7] Originally a generalized ideal demand system (GAIDS) was specified. However, preliminary analysis with this system provided unsatisfactory results, e.g., positive own-price elasticities for powdered-milk products. The purpose of this article is to study melamine impacts on different dairy food and not to test demand theory *per se*, thus the GAIDS model was replaced by a semilog model.

[8] The prices are interpolated from quarterly to monthly by CPI deflator.

[9] Data source is based on the representative sample of the 36 big and medium-sized cities of China.

[10] A data appendix, containing sources, is available upon request from the author.

monthly consumer’s expenditure divided by monthly per capita consumption, respectively. All of the expenditure variables were published in the same DAC source and China Statistical yearbook Database Report. Media coverage index variables for the melamine incident used in the analysis were also monthly data over the same period, constructed as discussed previously. Besides, the effects of seasonality using monthly demand binary variables and a linear time trending were incorporated in the model as well. Table 1 provides descriptive summary of the non-binary variables.

1.6 Empirical Results

In the empirical analysis, “dairy” is treated as a weakly separable group including liquid milk, powdered milk and yogurt in which consumption of an individual dairy item depends only on the expenditure of the group, the prices of the goods within the group, and certain introduced demand shifters. The preliminary tests demonstrate that there is no heteroscedasticity or autocorrelation in any of the equations. The demand equations were measured as a system by using Seemingly Unrelated Regression (SUR) under the assumption that error terms are correlated across equations but not over time. The results of the estimation are shown in Table 1.2.

Price and Expenditure Effect

The expenditure estimates of each milk type are all positive as expected (Table 1.2) and meanwhile, consistent with the law of demand, the own-price estimates are all negative and significant, among which liquid milk displays the greatest price sensitivity in the group.

As seen from Table 1.3, the expenditure elasticities for liquid milk, powdered milk and yogurt are 1.08, 0.50 and 1.71, respectively; except for powdered milk, two other products both generate statistically significant expenditure effects. Every one percent growth of yogurt consumption will lead to the greatest expenditure increase of dairy group. The most recent studies by Zheng et al. (2009) reported an average expenditure elasticity of 1.37 for dairy products in Jiangsu province, China that falls well within the reported range of 1.08 to 1.71.

The estimated own-price elasticities are –0.63 for liquid milk, –0.43 for powdered milk and –0.15 for yogurt, and apparently the own-price effect of liquid milk is the largest among dairy grouping. This figure is very close to the range of estimated fluid milk price elasticities of –0.66 to –0.73 reported by Kinnucan (1983) and the figure of –0.70 estimated by Smith et al. (1988). tabl(2010), milk was the most studied category aside from meat in United States. Thirteen studies provided elasticity estimates for specific milk fat levels. Mean elasticities for skim, 1%, and whole milk ranged from –0.75 to –0.79, whereas the mean elasticity for 2% milk was –1.22. Also, Bai et al. (2008) applied the Tobit model drawing on individual consumer survey data collected in urban Qingdao of China in 2005 and found that the own-price effect of fluid milk is –0.44, implying that fluid milk in Qingdao is a normal good but is price inelastic. However, unlike the previous studies that were based on either regional data or earlier time periods, the present analysis draws a recent national monthly dataset with respect to the melamine incident, therefore, allowing for the different estimated own-price effects.

The cross-price elasticity of liquid milk with respect to powdered milk price is 1.71, while the cross-price elasticity of powdered milk with respect to liquid milk price is only 0.23; which suggests that liquid milk is a very strong substitute for powdered milk, but powdered milk is quite a weak substitute for liquid milk.[11] This result, however, is consistent with the fact that much less powdered milk than liquid milk is consumed in urban China (Figure 1.1). And a plausible explanation for the asymmetry is that the dried milk price is ten times more than that for fresh milk given the same quantity, resulting in a larger income effect when the price of powdered milk increases, *ceteris paribus*, than when the price of liquid milk increases; also, it might be due in part to the melamine incident that made consumers instead turn to purchase the cheaper substitute of fresh milk.

Media Coverage Effect

The results show that incremental increases in current media coverage indices elicit small but negative impacts on the demand across the three products, but only powdered milk estimates support the theory that coefficients on both of current and lagged media indices should be negative. Similar to the Hawaii case, the estimated lagged effect of media indices for powdered milk exhibits a geometrically declining shape with the greatest impact occurring in the month of melamine news release, though the magnitude is far smaller than in the previous example. This yields the implication that within dairy grouping, direct demand response of powdered milk was most adversely affected, compared to that of fresh milk or yogurt, by the same media reports.

Likewise, the dummy variable estimate that distinguishes the *ex ante* and *ex post* incidence is negative for liquid and powdered milk, but not for yogurt, confirming the time trend of declined consumption of both liquid and powdered milk in the *ex post* period. The positive coefficients of three lags for yogurt indicate that the melamine crisis had no negative carryover effect on consumption of yogurt. Compared to the heptacholor incident in Hawaii decades ago,

[11] The cross-price effects for dairy products other than liquid milk and powdered milk were insignificant and therefore were dropped from the demand equations.

the melamine case occurring in China had much less impact on dairy demand. There might be several reasons; firstly, dairy products are a normal good as examined in the previous section and urban consumers of China can find very few suitable substitutes outside the dairy group, particularly those for infant food. In addition, the negative media coverage should be downward biased to some extent since in PR China the censorship mechanism for mass media is extremely restrictive in order to prevent public fears growing nationally; besides, even if the news media were reliably measured, temporary media coverage would be expected to have a weaker effect than the continuing warning (Schucker et al., 1983). Since the *Sanlu* dairy company was shutting down at the end of 2008, the credence value of milk might be regained from the "uncontaminated" or "qualified" milk produced by other dairy companies, including imported brands.

Seasonality Effect

The results also illustrate seasonality dummy and time-trending effects for the dairy consumptions; the yogurt demand had been increasing within the second quarter of the year and peaking in each August, while sales of both liquid milk and powdered milk were reducing in the summer time (third quarter) of year. But over the whole sample period, there was a gradually declining demand for both liquid and powdered milk, corresponding to the estimated effects of dummy variable mentioned above.

Demand Response Simulations

To further investigate the magnitudes of the media coverage impacts on each dairy demand, separate simulations for each of the own-demand responses were carried out using the estimated demand model while keeping media indices constantly zero over the *ex post* period; the results of the simulation shown in Table 1.4 indicate that in the period of 2009–10, the decline in annual demand of per capita powdered milk and yogurt would have been 9.09 and 0.47% less respectively without news coverage on the incident (holding all else equal), coinciding with significant increase of media indices during period of 2010; however, the decline in annual per capita liquid milk demand would have been 1.87% more. By contrast, the situation is different for the period of 2009 and 2011: without media coverage effects, the decline for liquid milk and yogurt consumption would be 0.08 and 4.67% less whereas the increase for powdered milk demand would be 5.71% less. These results, on the one hand, corroborate the empirical result mentioned earlier that liquid milk and powdered milk are substitutes and on the other, imply that the lagged effects of bad information on consumers' demand for dairy food were not sustained in the long term.

Meanwhile, the immediate demand response of specific month can be significantly larger than the reported averages over the sample period. For example, the difference in actual and projected own-demand response of powdered milk to media release was around –0.005 kg per person (a 9.14% decline) in Feb.2010 and –0.004 kg per person (a 9.10% decline) in Aug.2010, coinciding with the most largest effects of the media coverage indices in the two months. The average difference of actual and projected own-demand responses of powdered milk to media reports over the sample was only as much as –0.001 kg person per month. In sum, these deviations illustrate that the reported average demand response for powdered milk might be economically smaller than the immediate demand response to the influential (above 30) media indices; also, the empirical results suggest that as news coverage becomes sharply intense, the response to media effects would be larger, yet the apparent significant impacts are short-lived past the large increase in media coverage indices concerning dairy safety.

1.7 Concluding remarks

The focus and interest of this chapter is to investigate whether the news information concerning the *Sanlu* melamine incident has impacted the weakly separate grouping of dairy consumption in urban China over the recent years. The melamine incident exposed the government's lack of related legislation and supervision of dairy food safety, so that the powerful implementation of the Chinese Food Safety Law should be called for.

The economic and empirical framework reveals coefficients measuring the own- and cross-price effects of the dairy products in the case of melamine incident. The current media coverage effect indicated that publicized news information of melamine incident was detrimental toward demand for each of the studied products, while only powdered milk was showing consistent results with the hypothesis that the estimated lagged media effects were declining geometrically. The positive cross-price elasticity of liquid milk and powdered milk explains the reciprocal substitution effect and the asymmetric demand response to adverse media news.

The current estimated demand response to the melamine contamination incident over the study period was found to be economically small; there was also no significant evidence of negative cumulative effects (or lagged effects) on demand, especially in comparison to price effects and to previous estimates of unfavorable information effects regarding milk incidents. That the news media reports had no apparent significant impact on dairy sales, insofar as the statistical analysis was concerned, is not altogether unexpected. The sample of news indices was not national in scope covering urban China, whereas the monthly sales observations were. In contrast to the isolated island of Hawaii (the heptachlor case 1982) that was dominated by the domestic produced milk, the Chinese dairy industry has been seriously affected by the substantial increase of imports in last few years. The aggregate dataset did not distinguish domestic suppliers from imported producers, which otherwise might have yielded a different conclusion on the demand response to the incident.

Therefore, the media coverage of the melamine incident can be characterized as having a minor long-run impact on dairy demand accompanied with temporary but important shocks to demand.

Finally, the robustness of the results is subject to further scrutiny across alternative model specifications; and a series of extensions including using more refined data that measure media coverage indices or disaggregated cross-sectional household data on per capita consumption remains a topic for further investigations.

Chapter II. Examination of Asymmetry in Farm-Retail Price Transmission for U.S. Whole Milk

2.1 Background

In the recent decades, testing for Asymmetric Price Transmission (APT) and analyzing the APT elasticity are of importance in applied economics. However, a wide variety often conflicting theories of, and empirical tests for, APT co-exist in the literature. Economists who study market processes are therefore interested in price transmission processes. Of special interest are those processes that are referred to as asymmetric, i.e. for which transmission differs according to whether prices are increasing or decreasing (Meyer et al., 2004), versus others holding that retail prices would respond in the same manner for both increases and decreases in farm prices (Aguiar and Santana, 2002). In an extensive study of 282 products and product categories, including 120 agricultural and food products, Peltzman (2000) found asymmetric price transmission to be the rule rather than the exception, leading to the strong conclusion that the standard economic theory of markets is wrong, because it does not predict or explain the prevalence of asymmetric price adjustment; while others indicated that standard tests (such as the test applied by Peltzman) can result in excessive rejection of the null hypothesis of symmetry under common conditions.

2.2 Literature Review

Ward (1982) suggested that market power can lead to negative APT if oligopolists are reluctant to risk losing market share by increasing output prices. In a similar vein, Bailey and Brorsen (1989) considered firms facing a kinked demand curve that was either convex or concave to the origin. Otherwise if the firm conjectures that all firms will match an increase but none will match a price cut (convex), positive asymmetry will result.

Damania and Yang (1998) in a paper on imperfect information in a competitive duopoly stressed potential punishment as a cause of asymmetry. In their model demand was assumed to fluctuate randomly between high and low states. Punishment occured if a firm believes that its competitor is undermining a collusive price.

Besides, McCorriston et al. (2001) also showed that if an industry is characterized by non-constant marginal costs, there can be a significant impact on price transmission. The most important conclusion was that under certain conditions, price transmission may be greater in industries with increasing returns to scale (IRS) than in markets characterized by perfect competition and constant returns to scale (CRS).

Lloyd et al. (2009) investigated the buyer power in U.K. food retailing by a ‘first-pass’ test using a cointegrated vector autoregression and identified the null of perfect competition can be rejected in seven out of the nine food products, except for milk product no evidence can be found for the exercise of buyer power. Their proposed test, at the very least, offered a means via which the behavior of the retail-producer price spread is consistent with the buyer power.

However, some other economic analyses were focusing on price transmission elasticites and even on the demand/supply shifter impacts on dairy food. Kinnucan and Forker (1987) discussed the farm-retail price transmission effects and the work of distinguishing retail demand from farm supply shifts original derived by Gardener model. They identified that farm-retail price transmission process in the dairy sector was asymmetric using Houck procedure with data covering from 1971-1981 and tested the role in pricing asymmetry of retail demand versus farm supply shifts via a Chow-type test. Results suggested the major impact on retail prices of a change in the farm price of milk was felt sooner when farm prices are increasing than when farm prices are decreasing. Later Lass et al. (2000) estimated for Boston and Hartford retail dairy prices by using an econometric model, concluding that asymmetric speeds of adjustment to farm price increases and decreases were found and retail prices do return to the same level following equal farm price increases and decreases.

Recently Capps and Sherwell (2007) analyzed the behavior of tests for asymmetry by both of the conventional Houck approach and the error correction model (ECM) approach, employing monthly data over period from January 1994 to October 2002 for whole milk and 2% milk for seven typical U.S. cities. Their study was consistent with Kinnucan and Forker that price transmission elasticities were generally larger than corresponding elasticities associated with falling farm prices though they all were inelastic. By contrast, very few empirical tests yielded the opposite results that retail prices were more sensitive to decreases in farm prices than to increases in farm prices (Ward 1982, Punyawadee et al. 1991).

In this study, application of markup model and computation for the natural by-product of price transmission elasticities will be extended; meanwhile, testing the differences between demand as well as supply shifters will also be yielded with the *ad hoc* econometric procedures.

2.3 Theoretical Model

There are two alternative approaches in detecting asymmetric price transmission in this article and the emphasis is put on price transmission between the farm and retail levels of the vertical market system. But different from some previous analyses centered on spatial considerations by regional investigations (Bailey and Brorsen 1989, Frigon et al.1999, and Capps and Sherwell 2007), this paper is mainly based on asymmetric responses at the national aggregate level which might not lose any generality.

The "Houck" Approach

Houck (1977) developed a test for asymmetric price transmission in terms of the segmentation of price variables into increasing and decreasing phases. Till today, there are many analysts that have followed suit. The static asymmetric model can be written as:

(1) $\Delta P_{rt} = \alpha_0 + \alpha_1 \Delta P_{ft}^{+} + \alpha_2 \Delta P_{ft}^{-} + \varepsilon_t$

where P_{rt} and P_{ft} are retail and farm prices of the marketing chain, respectively, t=1,2,…,T, Δ is the first difference operator and following Houck procedure, $\Delta P_{ft}^{+} = P_{ft} - P_{ft-1}$ if $P_{ft} > P_{ft-1}$ and 0 otherwise, and $\Delta P_{ft}^{-} = P_{ft} - P_{ft-1}$ if $P_{ft} < P_{ft-1}$ and 0 otherwise. An example of the segmentation procedure is presented in table 2.1[12].

According to Houck (1977), Ward (1982) and Kinnucan and Forker (1987), the theoretical model including both asymmetry and lags is specified as follows:

(2) $\Delta P_{rt} = \pi_0\, TR + \sum_{i=0}^{m1} \pi_{1i} \Delta P_{ft}^{+} + \sum_{i=0}^{m2} \pi_{2i} \Delta P_{ft}^{-} + \pi_3\, MD_t + \varepsilon_t$

where ΔP_{rt} is the dependent variable; MD_t is the marketing cost difference variable, expressed as deviations from its initial value and *TR* is a trend term; at meanwhile π_{1i} and π_{2i} are coefficients in Equation 2 indicating the impact of rising and falling phases of farm whole milk prices on retail prices respectively; m1 and m2 represent the length of the lags with regards to rising farm prices and falling farm prices; ε_t is a random disturbance term.

$$FR_t = F_1 + \sum_{i=0}^{t-2} \max(\Delta P_{ft}^{+}, 0)$$

[12] Note that it makes no difference whether the variable is lagged and then segmented or segmented and then lagged.

measures the accumulated increases in farm prices up to period t,

$$FF_t = F_1 + \sum_{i=0}^{t-2} \min(\Delta P_{ft}^{-}, 0)$$

measures the accumulated decreases in farm prices up to period t. Theory provides no guidance for the number of lags to include. The model presented in the equation above is completely general, allowing for different lag lengths for rising and falling farm prices. Here I evaluated a number of different lag structures during the analysis and found that short-run price (period t) and two lagged prices (for periods t-1 and t-2) best fit the data for the Houck procedure.

Besides, the coefficients of π_1 and π_2 in equation (2) are representing the net effect of rising / falling farm prices on retail prices respectively. A formal test of linear restriction for the asymmetry hypothesis is below, which would be measured by t-test (Johnston, 1972).

H_N: $\pi_i^{+} = \pi_i^{-}$, for lags i=0, 1, 2;

and: H_N: $\sum_{i=0}^{m1} \pi_{1i} = \sum_{i=0}^{m2} \pi_{2i}$

The alternative hypotheses in each case are that the parameters or sums of parameters are not equal. The first hypothesis is a test of the speeds of adjustment of rising versus falling farm prices, sometimes referred to as a test for short-run price transmission asymmetry. For example, suppose the estimated parameter for current rising farm price is statistically greater than the estimated parameter for current falling farm price. Then processors will have a greater response to an increase in the current farm price than they will to a decrease in the current farm price. This result will demonstrate that upward adjustments in retail prices due to rising farm prices occur more rapidly than do downward adjustments due to falling farm prices (Lass et al., 2001). Correspondingly, the second hypothesis will test for long-run price transmission asymmetry which constitutes a linear combination of coefficients. A rejection of H_N is evidence of asymmetry or non-reversibility in price transmission. If one fails to reject H_N, then there exists evidence to support the notion of symmetry (or reversibility) in price transmission (Capps&Sherwell, 2007).

Meanwhile, I also computed mean lags for rising and falling farm price effects. And the formula is (Rao and Miller1971):

$$(3) \qquad ㄱL = \frac{\sum_{i=0}^{2} |\pi_i{}^{+/-}| * i}{\sum_{i=0}^{2} |\pi_i{}^{+/-}|}$$

where ㄱL is the mean lag for rising (falling) farm prices and the maximum lag length is two months. A mean lag for rising farm prices that is smaller than the mean lag for falling farm prices will identify that the upward speed of adjustment is rapid than the downward speed of adjustment.

The final set of measures of short-run and long-run elasticites are computed as follows:

$$(4) \qquad \varepsilon^{SR} = \pi_0^{+/-} * \frac{P_{ft}(\text{mean})}{P_{rt}(\text{mean})}$$

$$(5) \qquad \varepsilon^{LR} = \sum_{l=0}^{2} \pi_l^{+/-} * \frac{P_{ft-l}(\text{mean})}{P_{rt-l}(\text{mean})}$$

where all the estimated elasticities are all based on mean values of retail prices and farm prices. The *short-run* and *long-run* elasticities differ in that the latter incorporate all lagged farm price effects on the retail price. In this study, the current (*short-run*) period and one-month as well as two-month lags are used to calculate *long-run* elasticities.

The "Error Correction Model" Approach

The recent literature dealing with the price transmission has paid attentions to the time-series properties of the data, considering the inherent nonstationarity of prices or long-run stationary equilibria (cointegration) relationships among prices. Cramon-Taubel (1998) studied asymmetric price behavior in German producer and wholesale hog markets. Goodwin and Harper (2000) investigated linkages among farm, wholesale, and retail markets utilizing cointegration techniques in analyzing price transmission and asymmetric adjustment in the U.S. pork sector. Also, Goodwin and Piggott (2001) evaluated linkages among four corn and four soybean markets in North Carolina using cointegration methods. Till lately, Capps and Sherwell (2007) evaluated the behavior of price asymmetry for fluid milk according to both Houck procedure and error correction model for seven U.S. cities. Following Capps and Sherwell, this study is updated by the most recent monthly data and extended up to the national aggregate level.

Since the asymmetric ECM approach is motivated by the fact that none of the variants of the aforementioned Houck approach is consistent with cointegration between the retail and farm price series (Cramon-Taubadel, 1998; Cramon-Taubadel &Loy, 1999), the tests and discussion for ECM procedure is particularly of importance. If P_{ft} and P_{rt} are cointegrated, and the ECT can be developed into positive and negative components (Granger and Lee, 1989), consequently, the asymmetric error correction model in this analysis can be replicated from equation (7) by Capps and Sherwell (*op cit*) and written as:

(6) $$\Delta P_{rt} = \beta_0 \, TR + \sum_{i=0}^{p1} \beta_{1i} \Delta P_{ft}^{+} + \sum_{i=0}^{p2} \beta_{2i} \Delta P_{ft}^{-} + \sum_{i=1}^{p3} \beta_{3i} \Delta P_{rt-i} + \beta_4 ECT_{t-1}^{+} + \beta_5 ECT_{t-1}^{-} + \beta_6 MD_t + \varepsilon_t$$

Compared with Houck approach given by Equation (2), Equation (6) has three additional terms: $\sum_{i=1}^{p3} \beta_{3i} \Delta P_{rt-i}$, $\beta_4 ECT_{t-1}^{+}$ and $\beta_5 ECT_{t-1}^{-}$. Thus, the asymmetric ECM nests the Houck model when the lag lengths m1 and p1 or m2 and p2 are the same. Here in this paper, the lag structure of rising farm prices for ECM is set up differently with Houck procedure in terms of best-fitness principle. Therefore, to determine whether one model is superior to the other, the Akaike Information Criterion (AIC) (Akaike, 1974) or the Schwarz Information Criterion (SIC) (Schwarz, 1978) can be appealed in choosing between Houck and ECM specifications[13].

Absolute Marketing Margin vs. Relative Marketing Margin[14]

Revisited from Gardener's (1975) analysis, the summarized concluding was quoted as follows: "one implication of the results is that no simple markup pricing rule-a fixed percentage margin, a fixed absolute margin, or a combination of the two can in general accurately depict the relationship between the farm and retail price. This is so because these prices move together in different ways depending on whether the events that cause the movement arise from a shift in retail demand, farm supply, or the supply of marketing inputs." The three analytical expressions for the price transmission elasticites drawn from Gardner are as follows[15]:

$$(7)\quad \tau^{\mathrm{RD}} = \frac{\sigma + S_a e_b + S_b e_a}{\sigma + e_b}$$

$$(8)\quad \tau^{\mathrm{FS}} = \frac{S_a(\sigma + e_b)}{e_b + S_a\sigma - S_b\eta}$$

$$(9)\quad \tau^{\mathrm{MS}} = \frac{\sigma + e_a}{\sigma + \eta}$$

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[13] Choice of model selection rests on the lowest values of the AIC or SIC (Capps and Sherwell, 2007)

[14] Most of this part was replicated by the previous work of Kinnucan and Tadjion (2013).

[15] τ^i represents the percentage change in retail price over percentage change in farm price; η is the own-price of demand for the retail product x; σ is the elasticity of substitution between the farm-based input a and the bundle of marketing inputs b; e_a is the own-price elasticity of supply input a while e_b is the own-price elasticity of supply for input b; S_a the cost share for input a and S_b is the cost share for input b (S_b=1-S_a). The determinants of τ^i for isolated shifts in retail demand (RD), farm supply (FS) and marketing inputs' supply (MS).

To accommodate Gardner's analysis that the farm-retail price transmission elasticity in general will differ depending on the source of the supply or demand shock, we extend two equations from Gardener and Lloyd et al. (2009) testing for marketing margin stickiness with respect to farm price and price transmission effect that allows changes in whole milk quality to affect retail price.

$$(10)\Delta\ln M_t = \alpha_0 + \alpha_1 \Delta\ln P_{ft} + \alpha_2 \Delta\ln P_{ft} * D + \alpha_3 \Delta\ln FMC_t + \alpha_4 \Delta\ln INC_t$$

$$+\alpha_5 \Delta\ln CPRICE_t + \varepsilon_t$$

$$(11)\ \Delta lnP_{rt} = \beta_0 + \beta_1 \Delta lnP_{ft} + \beta_2 \Delta lnP_{ft} * D + \beta_3 \Delta lnFMC_t + \beta_4 \Delta lnINC_t +$$

$$\beta_5 \Delta lnCPRICE_t + \mu_t$$

where t=1, 2, 3…T indicates the time period of the observation; the delta terms represent first difference (e.g. $\Delta\ln M_t = \ln M_t - \ln M_{t-1}$), and M_t stands for the marketing margin, measured as the difference between the retail price and the retail farm price; here it is meaningful to distinguish between the *relative* marketing margin $M_R = P_{rt}/ P_{ft}$ and the *absolute* marketing margin $M_A = P_{rt} - \theta P_{ft}$ [16] since Gardner analyzed the relative margin, but not the absolute margin. D stands for the dummy variable that equals 1 while Farm price is falling, otherwise equals 0; FMC is the original food marketing cost index; INC is per capita real disposable personal income as proxy for retail demand shifters while $CPRICE$ is corn price which is a proxy for supply shifters[17]. The model is expressed in logarithmic first-difference form[18], with these

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[16] It is based on the retail-equivalent farm price $M = P_x - \vartheta P_a$ (P_x is the retail price and P_a is farm price) in terms of USDA's definition of the marketing margin and in the special case ϑ is a fixed constant (Gardener 1975, p.405; Reed *et al*.2002, p.2) where in this article, ϑ is assumed to be 1 in the dairy market. For absolute marketing margin, the logarithmic first difference may be approximated as: $\Delta\ln M_t \approx ((lnP_{rt} - lnP_{rt-1})$-$S_{1t}$ $(lnP_{ft} - lnP_{ft-1}))$, where S_{1t}= $((S_{1t}+ S_{1t-1})/2$ is the average farmers' share of the retail dollar between adjacent time periods.

[17] The supply shifter variable can also be represented by price index rather than corn price, such as cow slaughter price, manufacturing production costs etc., depending on the data access or availability.

variables of marketing cost, retail demand shifters and farm supply shifters all considered as exogenous[19]. And simply put, when it is legitimate to treat the price of marketing inputs exogenous, theory indicates that shifts in retail demand and farm supply have no effect on the *absolute* marketing margin; however, the both of demand and supply shifters should have statistically significant impact on the *relative* marketing margin. An intercept is included in the equation (10) to capture movements in the margin due to autonomous factors such as technical change in the marketers' production function.

According to Lloyd et al.'s test, when equations (7) and (8) are reduced to $\tau^{RD} = \tau^{FS} = S_a$, it can be implied farm supply and retail demand shifters have no effect on the marketing margin and price transmission elasticity in either case equals to the farmer's share of the consumer dollar. The two hypotheses tests are conducted with respect to equation (10) and (11):

Weak Form Hypothesis:

$$H_N: \alpha_4 = \alpha_5 = 0$$

H_A: H_N not true.

Strong Form Hypothesis:

$$H_N: \alpha_1 = \alpha_4 = \alpha_5 = 0 \text{ and } \beta_1 = \text{average Farmers' share}$$[20]

H_A: H_N not true.

In this instance, these combined zero-restrictions can be tested with a standard *Wald* and *F*-statistic.[21]

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[18] The logarithmic first differences are apt to be stationary.
[19] The Hausman test for marketing cost will be shown later.
[20] The average farmers' share of the retail dollar over the sample period.

2.4 Data and Estimation Procedures

Monthly undeflated (nominal) retail price of U.S. city average for whole milk, the undeflated farm price measured by announced cooperative price for whole milk, and monthly total food-marketing cost indices during the period January 1996—December 2011 from the national level were used in this study. The main Data source comes from Agricultural Marketing Service, U.S. Department of Agriculture (USDA) and Bureau of Labor Statistics (BLS)[22]. However, since the BLS does not appear to have retail price data for whole milk per gallon prior to 1995, and instead has price data per half gallon for 1980-95, changes occur over the years in how people buy milk (Half gallons used to be the most popular form. It is now gallons.) Thus, I select the time period for monthly and it lasts from January 1996 to December 2011, including 192 observations (the statistical summary is shown in table 2.2); and the farm and retail prices are expressed in terms of dollars per gallon.

I tested the hypothesis that farm prices Granger cause retail prices and vice versa. If farm prices Granger cause retail prices, then in the case where retail price is the dependent variable, the F-test corresponding to all coefficients associated with lagged farm prices should be statistically significant. Vice versa, if retail prices fail to Granger cause farm price, the F-test corresponding to all coefficients associated with lagged retail prices should not be statistically significant where farm price in the dependent variable (Granger, 1969). As shown in Table 2.3,

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[21] As emphasized by Lloyd *et al.* (2009, pp.4-5), although rejection of the zero restrictions constitutes evidence against perfect competition, failure to reject does not constitute evidence in favor of imperfect competition. Thus, failure to reject the null should not be taken as evidence that antitrust action is warranted. It is only when the null is rejected that such action might be appropriate. The qualification is necessary because other factors, such as non-constant returns to scale, could account for the rejection. In this sense, the tests are appropriately regarded as “first pass”, i.e., a signal that more in-depth (and costly) analysis is necessary to establish probable cause.

[22] Original data is available from the author upon request.

the Granger causality tests regards to the farm and retail prices of whole milk in this study support the underlying assumption that farm price precede or Granger cause retail prices, which means the retail price should be set up as the dependent variable, rather than farm price. This result hence coincides with the unidirectional-upward causal relationship assumption by Kinnucan and Forker (*op cit*) as well as with the same Granger causality test by Capps and Sherwell (*op cit*).

The next step was to test cointegration between the respective farm price and retail price series. The Augmented Dickey-Fuller (ADF) test was used to check on the stationarity of the retail and farm price series. And the Johansen test is also presented here to check for cointegration (see Table 2.4). Based on both the trace test and maximal eigenvalue test statistics, farm and retail prices of whole milk were cointegrated. Consequently, the asymmetric ECM can be applied to the respective cointegrated series compared with the traditional Houck model.

About 40% of the observations in the sample were months of farm price decline. And in terms of Kinnucan and Forker (*op cit*), this provided a sufficient number of price declines to reliably assess the asymmetry issue by the statistical procedures used here.

Besides, the changes in farm milk prices may not be matched by changes in retail prices for whole milk. Generally speaking, the changes between vertical prices became tend to track relatively closely over the sample period and seen from Figure 2.1, the retail prices do move when farm milk prices drop while it takes time for the shocks to pass through the firms that manufacture and distribute whole milk, which may contribute to the consideration of lag structures of equation (2) and (6). In short, differences in farm and retail prices as well as in farm-retail price spreads (market margin) in States are likely the results of government policy and the cost of transporting fluid milk from surplus to deficit areas concluded by Capps and Sherwell (2007).

Finite distributed lag structure in the farm price variable is assumed and in this article, the lag length for whole milk is two month except that only one month period for rising farm price by ECM procedure.

The length of lag distribution was based on Almon procedure (Almon, 1965), set up as a second order polynomial. Endpoint restrictions were used in conjunction with the Almon procedure. And the length of the distributed lag process was determined based on the Akaike Information Criterion (AIC) (Akaike, 1974) or the Schwarz Information Criterion (SIC) (Schwarz, 1978).

Ordinary Least Squares estimates are presented for both of Houck procedure and ECM approach since the serial correlation was not evident in those equations.

Furthermore, Seemingly Unrelated Regression (SUR) is applied for equation (10) and (11). And the quarterly data are estimated for the marketing margin model and price transmission model respectively, in that the demand and supply shifters will be paid more attentions, rather than focus on the length of lag distribution.

2.5 Empirical Results

The estimated coefficients and their standard errors generated by equation (2) and (6) are exhibited in Table 2.5-2.6. For the equations corresponding to the Houck approach, the goodness-of-fit statistics ranged from 0.60 to 0.64, and for the equations corresponding to the ECM approach, the goodness-of-fit statistics ranged from 0.70 to 0.71 which both suggest that the variation in retail price provided a reasonably good explanation, with or without the variables of total food-marketing costs and the trend term.

With the Houck approach, the number of lags associated with both rising and falling farm price variables was two, meaning the time for milk prices at the retail level to adjust to either increases or decrease in milk prices at the farm level was roughly the same, consistent with the result derived by Capps and Sherwell (2007). As observed in table 5, the direction of the price change matters for both the current period effect and the one-period lag effect. The positive calculated t-statistic indicated the current period rising coefficient was statistically greater than the current period falling coefficient, while the negative t-statistic for the one-period lag indicated the opposite. The two-period lag coefficients were not statistically significant[23].

Nevertheless, with the ECM approach(see table 2.6), the number of lags associated with rising farm prices was only one, relative to the lag number for falling-farm prices of two, which represents milk prices at the retail level might adjust faster to increases in milk prices at the farm level than to decreases; the positive calculated t-statistic also indicated that current period rising coefficient was statistically greater than the current period falling coefficient, and the negative t-statistic for the one-period lag was not statistically different.

With the Houck approach and the ECM approach, there is empirical evidence for all models that retail whole milk series adjustments to rising farm milk prices were much more rapid than adjustments to falling farm milk prices, implying the estimated coefficients in general were significant and agreed with a *priori* expectations. At the meantime, the cumulative (long-run)

[23] Specifically, for Houck procedure, the estimated coefficients of cumulative (long-run) effect indicate that Model IV yielded a smaller number for rising-farm prices but a larger one for falling-farm prices; while for ECM, it did not show such a variance among different models.

effect on retail milk prices attributable to increases in farm milk prices exceeded the cumulative effect attributable to decreases in farm milk prices. The F-test associated with the null hypothesis that retail prices responded symmetrically to increases and decreases in farm prices was rejected with both approaches in all models.

Besides, to decide whether the error correction model is statistically superior to the Houck procedure, one may use either the Akaike Information Criterion (AIC) (Akaike, 1974) or the Schwarz Information Criterion (SIC) (Schwarz, 1978) to make comparison since the lag structures are not the same. Basically, each model of ECM approach provided a lower value of either AIC or SIC than Houck approach did and it might suggest the ECM be preferred over the Houck model based on model selection criteria, and statistically speaking, cointegration played a relatively vital role.

Furthermore, the computed Farm-Retail price transmission elasticities of short-run and long-run are displayed for both Houck approach and ECM approach[24] (Table 2.7), illustrating the unequal retail response to changes in the farm price of whole milk. All estimated elasticities of price transmission were quite inelastic. For rising-farm price transmission the long-run elasticities were similar while for falling-farm price transmission the long-run effects were almost twice as large as the corresponding short-run effects. On the other hand, the long-run rising price elasticity was more than the corresponding falling-price elasticity by *10.9* times for Houck model and *13.5* times for ECM procedure which demonstrates that increases in the farm price of whole milk were passed through to the retail level more completely than were farm price decreases; moreover, the mean lags associated with the rising farm price variables were uniformly smaller than the corresponding mean lags of the falling farm price variables, stating

[24] The elasticites are computed based on the estimated coefficients of Model III by both approaches, in order to compare with the estimates derived by Capps and Sherwell (2007).

that retail dairy product prices adjust much faster to increases in the farm price than to decreases. To compare the empirical results with those addressed by previous studies e.g. Kinnucan and Forker (1987)[25], who reported the elasticity of price transmission for rising farm prices of milk to be 0.274 in the short run and 0.462 in the long run, and for falling farm prices of milk the short-run estimate to be 0.184 versus 0.330 for long run, it can be perceived in a clearer fashion that our elasticities for falling-farm prices were much smaller than theirs but the long-run estimates for rising-farm prices were more close to theirs.

Lass et al. (2001) estimated northeast dairy compact impacts for Boston and Hartford retail prices and figured out that the immediate effect of a farm price increase was greater than the immediate effect of a farm price decrease while over time the aggregate effects of increases and decreases were approximately the same. Their estimates for rising-farm prices varied around from 0.30 to 0.46 and from 0.14 to 0.35 for falling-farm prices; though our results seemed outside their intervals (the rising prices are larger while the falling prices are smaller), there is one thing in common: that in our study the ECM approach derived the long-run rising elasticity of 0.429 which was a bit smaller than the short-run rising elasticity of 0.433; likewise, Lass et al. found for Boston market the long-run rising estimate reached 0.351 which was also smaller relative to its current (short-run) effect of 0.457[26]. Thus, it cannot be inferred that the retail response to long-run farm prices be slower to short-run prices.

Capps and Sherwell (2007), applying the same two procedures as this study did, illustrated the elasticities of price transmission for seven U.S. cities respectively. The numbers for rising farm prices of milk ranged from 0.037 to 0.263 in the short run and from 0.187 to 0.527 in the

[25] The long-run rising price elasticities exceed corresponding its counterpart by 40% for fluid milk, covering from 1971-1981.

[26] The short-run rising elasticity exceeds the long-run effect around 30% for Boston (Lass et al. 2001).

long run; and for falling farm prices of milk the elasticities varied from 0.005 to 0.166 in the short run and from 0.031 to 0.553 in the long run. Accordingly, with the same two approaches mentioned above, our results were more consistent with those of Capps and Sherwell, and agreed to the estimation evidence that increases in the farm price of milk were passed through to the retail level more fully than were decreases in the farm price of milk.

Potential importance for price transmission work of distinguishing retail demand from farm supply shifts was focused by some numerical results derived from the Gardener model (Table 2.8). And the relevance between Gardener model and equation (2) or (6) is due to the aforementioned Granger causality test that only farm prices have unidirectional impact on retail prices, rather than vice versa.

According to Hausman test results from table 2.9, t-tests were both failed to reject on 10% level for either equation (10) or (11)[27], indicating that the null hypothesis of exogeneity was compatible with the data with almost no probability of a type II error, which means the price of marketing inputs is exogenous (e_b=∞), as was commonly assumed in empirical studies (e.g., Heien 1980, Kinnucan and Forker 1987, Wohlgenant and Mullen 1987).

Estimated results were derived by SUR from Table 2.10. For U.S. whole milk in the sample period, and none of the equations showed signs of serial correlation (see *D.W.* test); one per cent increase of farm price for whole milk would yield to around 0.56% decreases for *relative* marketing margin and 0.44% increases for retail price, which identified the estimated price transmission elasticity for rising-farm prices was around 0.44 (t-ratio=11.10); and this

[27] The exogeneigty test was performed on *FMC* using per capita personal income, Producer Production Index, and Real Retail and Food Services Sales (millions of dollars), data available upon request.

result highly agreed to the long-run price transmission elasticity for rising-farm prices[28] of 0.43 with the ECM procedure. The demand and supply shifters had no any significant effect on neither of *absolute* marketing margin nor *relative* marketing margin, which satisfied one of a *priori* expectations that shifts in retail demand and farm supply have no effect on the *absolute* marketing margin, but did not agree to the other one that shifts should significantly influence on the *relative* marketing margin. Thus, it is hardly to detect if the non-competitive market prevails or not at this stage. Additionally, the marketing cost variable was statistically significant for the full absolute marketing margin model (i) as well as for the full price transmission model (v), where their estimated coefficients were close to each other (0.57 vs. 0.54). The technical change variable (*constant* in table 10), however, yielded the minus results across all the models which can be inferred the marketing margin be negatively affected by the technical change but limited to a minor extent.

Of key interest are the results relating to the demand and supply shifters, and test statistics were distributed as χ^2 under the null hypothesis of no buyer power (i.e. perfect competition). However, following Lloyd et al.'s (2009) methodology, the U.S. whole milk was found not to reject the perfectly competitive nulls, and seen from table 2.11, there was no evidence for the exercise of buyer power in milk. The two big players, Dallas-based Dean Foods Co. and the Dairy Farmers of America, a cooperative out of Kansas City, Missouri, were the target of pending federal class actions, one filed in 2008, and the other in 2010, in which they were accused of colluding to control market access and suppress milk prices (USDA, 2010)[29]. Thus if

[28] The estimated marketing margin and price transmission regression are both based on the quarterly data, which test only for the long-run effects.

[29] The source can be found by this link: http://www.bloomberg.com/news/2010-05-27/obama-regulators-to-review-dairy-farmers-complaints-of-market-dominance.html

the milk price spread was being maintained by collusion rather than competition, as the regulatory authorities have found to be the case, it is little wonder that our simple test was unable to detect what amounts to relatively sophisticated strategic pricing behavior.

Note also from Figure 2.2, either for the absolute market margin or for the relative margin, there was no overall trend in the retail-farm spread. And given the raw data, it is not surprising that the null hypothesis in this case cannot be rejected. There could of course be some other aspect of buyer/seller power that exists in this market and that the concerns about collusion between retailers and processors did not negatively impact on milk producers taken over the period for which our data applies (Lloyd et al.2009).

More specifically, for *absolute* marketing margin, the *Wald* statistic test regarding to both Weak-Form and Strong-Form hypothesis was too small to reject the null at any confidence level (Table 2.12). Thus, it resulted in the conclusion that retail demand and farm supply shifters for *absolute* marketing margin had no effect on the marketing margins for whole milk. Nevertheless, for *relative* marketing margin the Strong-Form of whole milk appeared quite differently, and the *Wald* statistic test rejected the null at any confidence level (Table 2.13). The computed chi-square value for the *Wald* test was 6880.97 and large enough to reject null hypothesis at any *p* value. This, corroborated with the *relative* marketing margin elasticity of –0.557% (t-ratio= –13.96) (An isolated 1% increase in the farm price of whole milk is likely to narrow the marketing margin by 0.56%). From this point, it is difficult to reveal any clear-cut information regarding the effects of market power on the degree of price transmission in the U.S. marketing channel for whole milk.

2.6 Concluding Comments

This study rests on the assumption that the aggregate technology for food processing and marketing is characterized by CRTS (Wohlgenant 1989, hold for all commodities except fresh fruits).

The larger size of rising farm price elasticities compared to decreasing price elasticities demonstrates that the retail whole milk prices adjust faster to rising farm prices than to decreases. And the slower response of retail prices to downward movements in farm prices stays consistent with the commonly held belief that consumers do not benefit from decreases in farm prices in dairy market.

Meanwhile, though it is inherently likely to constitute the evidence of non-competitive pricing (market power at the retail level) in the U.S. dairy/whole milk marketing channel driven by the presence of asymmetric price transmission mentioned earlier, the acceptance of the perfectly competitive null is apparently contradicting to the evidence for the exercise of buyer power in milk. In combination with the results provided by both *absolute* marketing margin and *relative* marketing margin, no meaningful conclusions can be drawn regarding the effects of market power on the degree of price transmission for U.S. whole milk[30].

It is probably that competitive market clearing does not require that the elasticity of farm-retail price transmission equal 1, in that Gardner (1975) pointed out situations in which this would occur are rare to the point of irrelevance. By the same token, price transmission

[30] Weldegebriel H.T. (2004) casted doubt on the validity of empirical estimates of the price transmission coefficient obtained from regressions using retail and farm level price information, quoted as "without knowledge of the functional forms of retail demand and farm input supply, little can be inferred… in particular, it is not generally possible to attribute low (or high) values of the price transmission coefficient to market power."

elasticities close to zero might represent non-competitive pricing, or they simply might represent consumers' preference for food products that intensive in the marketing input. From an econometric perspective, a price transmission elasticity close to zero in models that exclude marketing costs (Model III) perhaps is to be expected, as such models are subject to attenuation bias whenever consumers can substitute more easily than intermediaries, i.e., whenever $|\eta| > \sigma$.

The main contribution of this paper is to examine the asymmetric price transmission from multiple perspectives. Houck and ECM approaches are both employed and compared for this analysis, which corroborates the previous assumed unidirectional-upward causal relationship from farm to retail prices and updated Capps and Sherwell (2007) 's analysis from regional areas to the entire nation with more recent data. Besides, the comparison between *absolute* marketing margin and *relative* marketing margin are also shedding a new light that it is not generally possible to attribute low (or high) values of the price transmission coefficient to market power, or if the collusion rather than competition exists in U.S. dairy market, what amounts to relatively sophisticated strategic pricing behavior would be unable to detect.

There may be other necessary factors affecting marketing margin that include price risk, technical change and other structural change, product quality and seasonality (Wohlgenant, 2001). Since it is not accessible to Four-firm concentration ratios (CR4) in the milk processing sector, measuring the market power with approximated proxy variables remains a further topic.

Chapter III. Effects of Recession and Dollar Weakening on the U.S. Agricultural Trade Balance

3.1 Introduction

The J-curve hypothesis generated a series of empirical research that investigated the existence of J-curve both in US data and other countries' data. Earlier studies like Krugman and Baldwin (1987) found evidence of a J-curve in the US data, and Carter and Pick (1989) indicated the first segment of the J-curve did exist for the U.S. agricultural trade balance, based on the empirical evidence that a 10 percent depreciation of the U.S. dollar was estimated to lead a deterioration of the agricultural trade balance that would last for about nine months. However in a series of papers Rose and Yellen (1989), Rose (1990) and (1991), not only the J-curve hypothesis was rejected, but also it is argued that there was no significant effect of the real exchange rate on the trade balance for both the developing and the developed countries, including the U.S.

By the same token, Bahmani-Oskoee and Brooks (1999) used the ARDL approach to analyze the US data and found that short-run results supported Rose and Yen (1989) that there was no effect of real exchange rate on the trade balance in the short run, but in the long-run the real depreciation of the US dollar was found to have a favorable effect on the trade balance. Wilson and Tat (2001) on the other hand again by using the Rose and Yellen's model found similar results for Singapore. However Singh (2002) by using a trade balance model *a la* Rose (1991) and an error correction model implied that trade balance of India was sensitive to the real exchange rate changes as opposed to Rose (1990) and (1991). Akbostanci (2002) investigated the existence of a J-curve in the Turkish data in the period of 1987-2000 and suggested the results did not exactly support the J-curve hypothesis in the short-run, yet the short-run behavior of the trade balance in response to real exchange rate shocks showed an S-pattern reminiscent of the Backus et al (1994) rather than the J-curve pattern.

The most recent work of Baek et al. (2009) analyzed the dynamic effects of changes in change rates on bilateral trade of agricultural products between the U.S. and its 15 major trading partners, by applying the ARDL model of the error correction version, which, in the empirical specification is *ad hoc*; they concluded the exchange rate plays a crucial role in determining the short- and long-run behavior of U.S. agricultural trade, but there was no evidence of the J-curve phenomenon for U.S. agricultural products with its major trading partners. However, Baek et al. applied the bilateral trade balance model between U.S. and its 15 major trading partners, which did not specialize the difference among them, since these trading partners include developing as well as developed countries and their historical trading pattern with U.S. could be restricted to internal circumstances such as macroeconomic environment, political changes and national productions etc. Meanwhile, the exchange rate effects on U.S. agricultural trade should be distinguished between these selected countries because different countries have different policies for adjusting exchange rates. Therefore, in our study, we will do the aggregation analysis for U.S. trade balance and select exchange rate index based on world US agricultural trade weighted real rate (where year 2005=100), rather than on the real annual country exchange rates of Baek et al.; and our article will take the recession effects into account, the trade issue becomes different from the macro-economic perspective.

There have been numerous theoretical arguments and empirical analyses in the past decades, and exchange rate and the income growth particularly draw the most attention, yet there is no consistent conclusions to the dynamic effects on U.S. agricultural trade. The general objective of this article is to determine the U.S. agricultural trade pattern during the past decades. The two specific objectives: (a) determine if the exchange rate dominates the factors that impact on agricultural trade balance, and (b) determine if there is any recession effects on U.S. agricultural trade.

Unlike the previous work by Baek and Koo (2011), we divide U.S. agricultural products into two categories: bulk and high-value goods[31]; in addition, the four times decennial recessions from 1970s to date will be included and examined if each recession had different impacts on U.S. trade of agricultural products. Hence, the two individual commodities plus the combined products will be analyzed for their trade balance between United States and the Rest of World (ROW). Within our sample period, the U.S. trade balance of high-value and combined agricultural goods showed similar pattern especially after 1990s, and both were almost flat since 2008; while the balance of high-valued commodities had been in deficit mostly[32]. Bulk commodities, on the other hand, fluctuated greatly with ups and downs through each time of recession (see Figure 3.1).

During our sample period, the major oil shock occurred in the early 1980s which is also referred to as a "double-dip" or "W-shaped" recession led to a rising real U.S. interest rate, a decline in import demand, and stagnant growth in many external debt impacted developing countries (see Table 3.1). And prior to the early 1990s recession, both a devalued dollar and GDP growth helped to rise the U.S. agricultural trade balance, including individual bulk and high-value goods. The Economic Research Service (ERS) reported that since 1995, the exchange rate for U.S. bulk exports was up by nearly twenty percent. The Asian financial crisis that began in July 1997 led to depreciated currencies, decreased economic growth, and depressing global commodity prices, which decreased U.S. agricultural exports. And U.S. agricultural exports

[31] Instead of the three groups of bulk, intermediate and consumer-oriented products categorized by Baek and Koo (2011).

[32] Except in the period from 1992 to 1996, the trade balance of HVP exceeded than 1.

value experienced a 23 percent decrease in real terms for the period from fiscal 1997 to fiscal 1999. USDA analysis blamed oversupplies for this decrease, however, the U.S. as a non-crisis exporter experienced a four percent increase in economic growth, decreased producer prices, increased production, increased consumption, decreased exports and increased imports due to the Asian crisis, which included significant depreciations of crisis countries' currencies.

Over 2000–07, the per capita GDP of U.S. grew by around 1.5% a year, falling below the annual growth of 2.3% during the 1990s; meanwhile, the per capita GDP of the rest of world grew by 2.37% yearly from 2000-07, which was far more than the percentage of 0.88% since 1990s. The spread of the crisis beyond the United States is impacting economic growth throughout the world; the world GDP in 2009 dropped by 4.3%, compared to about 2.5% growth in 2008 and 3% yearly average growth since 1970 (see Figure 3.2). Our question is how might the decennial economic crisis and ensuing economic downturn affect U.S. agriculture goods over the four decades, especially given how important exports and imports are to the sector?

The economic crisis can have direct and indirect effects on U.S. agriculture. The direct effects come from changes within the U.S. economy alone. The indirect effects will occur from how the crisis impacts foreign income and trade and world energy prices (Liefert and Shane 2009). Seen from Figure 3.2, before each economic recession mounted there went through a devaluation of U.S. dollar, no matter how long it lasts; between 2002 and 2008 the dollar fell in real terms against all foreign currencies by 18%, 22% and 21% for bulk, high-value and combined agricultural products respectively. However, the crisis' short-run (current) effect on the U.S. dollar has been, perhaps ironically to appreciate rather than depreciate, yet the appreciation did not last long and reappeared in 2011 again. According to Liefert and Shane

(2009), the dollar was appreciating against the currencies of most other countries, developed as well as developing, which means the appreciation could hurt U.S. agricultural exports by making them less price competitive compared to output produced not only by importing countries but also by export competitors, such as Canada, Australia, and Brazil.

The remainder of the article is organized as follows: the next section introduces the theoretical foundation for the J-curve pattern of U.S. agricultural trade balance. Then, the empirical model will be described and the data set used will be discussed, followed by results analysis for both short- and long-run effects on U.S. agricultural trade. And the final section concludes the paper.

3.2 Theoretical Framework[33]

The basic economics of the agricultural trade balance in a partial-equilibrium setting can be illustrated with the aid of the following structural model. In this model, agricultural products are assumed to be undifferentiated in international trade, and we abstract from complicating factors such as price wedges due to subsidies and tariffs, transportation costs, and farm programs. Prices are assumed to be determined under competitive conditions so that the Law of One Price (LOP) holds. Exports and imports are a function of local currency prices, which adjust in response to changes in the exogenously-determined exchange rate and income levels in the importing and exporting countries. Other exogenous variables that shift supply and demand curves are suppressed. With these assumptions, the structural model consists of seven equations:

(1) $$Q_d = D(P_d, Y_d)$$

[33] This section was mainly written by Henry W. Kinnucan.

(2) $$Q_x = X(P_f, Y_f)$$

(3) $$Q_s = S(P_d)$$

(4) $$Q_m = M(P_f)$$

(5) $$P_d = P_f \cdot e$$

(6) $$Q_s + Q_m = Q_d + Q_x$$

(7) $$TB = \frac{P_f Q_x}{P_d Q_m} = \frac{P_d Q_x}{P_d Q_m e}.$$

Quantities of agricultural products produced and consumed in the home market in the initial equilibrium are given by Q_s and Q_d; quantities imported into and exported from the home market are given by Q_m and Q_x; the price of the agricultural products in foreign currency units (FCU) in the initial equilibrium is given by P_f, and the corresponding price in domestic currency units (DCU) is given by P_d; income levels in the domestic and export (foreign) markets in the initial equilibrium are given by Y_d and Y_f; the exchange rate in the initial equilibrium is given by $e = DCU/FCU$; and the trade balance in the initial equilibrium is given by *TB*. Because e tells how many units of domestic currency can be purchased per unit of foreign currency, an increase e implies currency weakening (devaluation) from the domestic buyers' perspective.

The model consists of seven endogenous variables (Q_d, Q_x, Q_s, Q_m, P_d, P_f, TB) and three exogenous variables (Y_d, Y_f, e). At issue are the hypotheses that can be deduced given the structure defined by equations (1) – (7). Specifically, do the reduced-form elasticities of trade balance with respect to income and exchange rate have determinate signs?

To determine that, the model is first expressed in proportionate change form as follows

(8) $$Q_d^* = \eta_d P_d^* + \gamma_d Y_d^*$$

(9) $$Q_x^* = \eta_x P_f^* + \gamma_x Y_f^*$$

(10) $$Q_s^* = \varepsilon_d P_d^*$$

(11) $$Q_m^* = \varepsilon_m P_f^*$$

(12) $$P_d^* = P_f^* + e^*$$

(13) $$k_s Q_s^* + k_m Q_m^* = k_d Q_d^* + k_x Q_x^*$$

(14) $$TB^* = Q_x^* - Q_m^* - e^*.$$

The asterisk (*) indicates proportionate change (e.g., $Q_d^* = dQ_d/Q_d$); the Greek symbols indicate elasticities; and the k terms indicate quantity shares. Specifically, η_d (< 0) and η_x (< 0) are price elasticities of domestic and export demand; γ_d (> 0) and γ_x (> 0) are income elasticities of domestic and export demand; ε_d (> 0) and $\varepsilon_m > 0$ are price elasticities of domestic and import supply; $k_s = Q_s/(Q_s + Q_m)$ is the share of domestic supply that comes from domestic production; $k_m = Q_m/(Q_s + Q_m)$ is the share of domestic supply that is imported; $k_d = Q_d/(Q_s + Q_m)$ is the share of domestic supply that is consumed in the home market; and $k_x = Q_x/(Q_s + Q_m)$ is the share of domestic supply that is exported.

Reduced Form

The reduced-form equations for proportionate changes in domestic price, exports, imports, and trade balance implied by the structural model are as follows:

(15) $$P_d^* = \left(\frac{k_m\varepsilon_m - k_x\eta_x}{\varepsilon - \eta}\right) e^* + \left(\frac{k_d\gamma_d}{\varepsilon - \eta}\right) Y_d^* + \left(\frac{k_x\gamma_x}{\varepsilon - \eta}\right) Y_f^*$$

(16) $Q_x^* = \left(\frac{-\eta_x(k_s\varepsilon_d - k_d\eta_d)}{\varepsilon-\eta}\right)e^* + \left(\frac{\eta_x k_d \gamma_d}{\varepsilon-\eta}\right)Y_d^* + \left(\frac{(\varepsilon - k_d\eta_d)\gamma_x}{\varepsilon-\eta}\right)Y_f^*$

(17) $Q_m^* = \left(\frac{-\varepsilon_m(k_s\varepsilon_d - k_d\eta_d)}{\varepsilon-\eta}\right)e^* + \left(\frac{\varepsilon_m k_d \gamma_d}{\varepsilon-\eta}\right)Y_d^* + \left(\frac{\varepsilon_m k_x \gamma_x}{\varepsilon-\eta}\right)Y_f^*$

(18) $TB^* = \left(\frac{(\varepsilon_m - \eta_x)(k_s\varepsilon_d - k_d\eta_d)}{\varepsilon-\eta} - 1\right)e^* + \left(\frac{(\eta_x - \varepsilon_m)k_d\gamma_d}{\varepsilon-\eta}\right)Y_d^* + \left(\frac{\varepsilon_m(k_m - k_x) + k_s\varepsilon_d - k_d\eta_d}{\varepsilon-\eta}\right)Y_f^*$

where $\varepsilon = (k_s\varepsilon_d + k_m\varepsilon_m) > 0$ is the overall supply elasticity and $\eta = (k_d\eta_d + k_x\eta_x) < 0$ is the overall demand elasticity. The reduced-form elasticities in equations (15) – (18) shed light on empirical estimates in the literature. For example, the only variable to have a determinant effect on trade balance is domestic income. Specifically, market theory predicts that an isolated increase in domestic income will reduce the trade balance, i.e., cause import value to rise in relation to export value. No such predictions are forthcoming about isolated movements in either the exchange rate or foreign income. Thus, one should not be surprised to find mixed signs (or a zero effect) for these variables in an empirical trade balance relation.

Exports are positively related to domestic currency devaluation and foreign income, and negatively related to domestic income. Imports are negatively related to domestic currency devaluation, and positively related to domestic and foreign income. In the trade balance literature these relations are typically estimated with the dependent variable defined as trade value (e.g., Baek and Koo, 2011). The relevant reduced-form equations in this instance are obtained by combining equation (15) and with equations (16) and (17) to yield:

(19) $V_x^* = \left(\frac{-\eta_x(k_s\varepsilon_d - k_d\eta_d + k_x) + k_m\varepsilon_m}{\varepsilon-\eta}\right)e^* + \left(\frac{(1+\eta_x)k_d\gamma_d}{\varepsilon-\eta}\right)Y_d^* + \left(\frac{(\varepsilon - k_d\eta_d + k_x)\gamma_x}{\varepsilon-\eta}\right)Y_f^*$

(20) $V_m^* = \left(\frac{-\varepsilon_m(k_s\varepsilon_d - k_d\eta_d - k_m) - k_x\eta_x}{\varepsilon-\eta}\right)e^* + \left(\frac{(1+\varepsilon_m)k_d\gamma_d}{\varepsilon-\eta}\right)Y_d^* + \left(\frac{(1+\varepsilon_m)k_x\gamma_x}{\varepsilon-\eta}\right)Y_f^*$

Switching from quantity to value renders the domestic income effect in the export equation and the exchange rate effect in the import equation indeterminate. An increase in domestic income will reduce export value only if export demand must be price elastic. Similarly, domestic currency devaluation will reduce import value only if the negative quantity effect of the devaluation outweighs the positive price effect. That is, referring to the coefficient of e^* in equation (20), only if $|-\varepsilon_m(k_s\varepsilon_d - k_d\eta_d)| > (k_m\varepsilon_m - k_x\eta_x)$.

Small Trader Effects

The analysis thus far indicates that the only variable to have a determinant effect on trade balance is domestic income. This suggests restrictions need to be placed on model parameters if market theory is to have predictive content with respect to exchange rate and foreign income. One such restriction is that the country in question is too small a buyer or seller in the world market to affect prices. A small seller in world markets implies the country faces a perfectly elastic export demand curve. In this instance, the equations (15) – (20) reduce to:

<u>Reduced form when $\eta_x = -\infty$ (small exporter):</u>

(21) $P_d^* = V_m^* = e^*$

(22) $Q_m^* = 0$

(23) $V_x^* = \left(\frac{k_s\varepsilon_d - k_d\eta_d + k_x}{k_x}\right)e^* - \left(\frac{k_d\gamma_d}{k_x}\right)Y_d^*$

(24) $TB^* = Q_x^* = \left(\frac{k_s\varepsilon_d - k_d\eta_d}{k_x}\right)e^* - \left(\frac{k_d\gamma_d}{k_x}\right)Y_d^*$

Imposing the small exporter assumption eliminates the ambiguity about the effects of exchange rate and foreign income on trade balance. Specifically, theory now predicts that changes in foreign income will have no effect on trade balance, and that domestic currency devaluation will improve the trade balance. The reason for the latter is that although devaluation causes import value (cost) to increase (since the full burden of the exchange rate adjustment falls on domestic price), export value rises faster. Specifically, $(V_x^*/e^* \,/\, V_m^*/e^*) = \left(\frac{k_s\varepsilon_d - k_d\eta_d + k_x}{k_x}\right) > 1$.

A small buyer in world markets faces a perfectly elastic import supply curve. In this instance, equations (15) – (20) reduce to:

<u>Reduced form when $\varepsilon_m = \infty$ (small importer):</u>

(25) $P_d^* = e^*$

(26) $Q_x^* = 0$

(27) $Q_m^* = -\left(\frac{k_s\varepsilon_d - k_d\eta_d}{k_m}\right)e^* + \left(\frac{k_d\gamma_d}{k_m}\right)Y_d^* + \left(\frac{k_x\gamma_x}{k_m}\right)Y_f^*$

(28) $V_x^* = e^* + \gamma_x Y_f^*$

(29) $V_m^* = -\left(\frac{k_s\varepsilon_d - k_d\eta_d - k_m}{k_m}\right)e^* + \left(\frac{k_d\gamma_d}{k_m}\right)Y_d^* + \left(\frac{k_x\gamma_x}{k_m}\right)Y_f^*$

(30) $TB^* = \left(\frac{k_s\varepsilon_d - k_d\eta_d}{k_m}\right)e^* - \left(\frac{k_d\gamma_d}{k_m}\right)Y_d^* + \left(\frac{k_m - k_x}{k_m}\right)Y_f^*$

Imposing the small importer assumption eliminates the ambiguity about the effects of exchange rate on trade balance, but not the ambiguity about the effects of changes in foreign income. The latter effect is positive only if the import share exceeds the export share. As for the exchange rate effect, domestic currency devaluation causes the trade balance to improve, the same result obtained for the small exporter case.

The J-Curve

The *J*-curve describes a situation where the trade balance initially deteriorates following devaluation before it improves. The Marshall-Lerner Condition attributes this phenomenon to differences in short- and long-run elasticities of demand (e.g., Rose and Yellen 1989; Rose 1991). The MLC in terms of the present model can be derived by dividing the trade-balance relation (equation (14)) through by e^* to yield

(31) $$\frac{TB^*}{e^*} = \frac{Q_x^*}{P_f^*}\frac{P_f^*}{e^*} - \frac{Q_m^*}{P_d^*}\frac{P_d^*}{e^*} - 1$$

where $\frac{Q_x^*}{P_f^*} = \eta_x$ is the export demand elasticity; $\frac{Q_m^*}{P_d^*} = \eta_m$ the import demand elasticity; $\frac{P_d^*}{e^*} = PTE$ is the pass-through elasticity into the domestic price; and $\frac{P_f^*}{e^*} = (PTE - 1)$ is the pass-through elasticity into the foreign price. Substituting these parameters into equation (21), for the trade balance to improve following devaluation the following condition must hold

(32) $$PTE\ |\eta_m| + (1 - PTE)|\eta_x| > 1.$$

where

(33) $$0 < PTE = \left(\frac{k_m\varepsilon_m - k_x\eta_x}{\varepsilon - \eta}\right) \leq 1.$$

Equation (32) is the MLC modified to include the pass-through elasticity. If the sum of the import and export demand elasticities (in absolute value) is less than one, as might be true in a short-run situation, the trade balance deteriorates following devaluation. This explains the downward-sloping portion of the *J*-curve. If pass-through is complete ($PTE = 1$), as would be true if export demand or import supply were perfectly elastic (the small-trader situation), the MLC reduces to $|\eta_m| > 1$. In this instance, the trade balance improves following devaluation provided import demand is price elastic. The reason is that with elastic import demand, the proportionate increase in price associated with devaluation is more than offset by a proportionate decrease in import quantity, implying a decrease in import value. Because foreign price and export revenue are unaffected when pass-through into the domestic price is complete, the trade balance necessarily improves.[34]

In summary, our structural model yields only one hypothesis, namely that domestic income growth will cause the agricultural trade balance to deteriorate. The trade balance deteriorates because an increase in domestic income causes a simultaneous increase in imports and decrease in exports (and price effects cancel owing to LOP). The effect of an increase in foreign income on the trade balance is indeterminate, as is the effect of domestic currency devaluation. Thus, the effect of these variables on the trade balance is an empirical issue.

3.3 Empirical Model

To test the hypothesis that the domestic income growth has a negative effect on the U.S. agricultural trade balance, we follow Baek *et al*. (2009) and replicate their work with employing an autoregressive distributed lag (ARDL) modeling approach. The ARDL approach avoids spurious regression associated with non-stationary time series, and permits distinguishing short-run from long-run effects. The ARDL approach lends itself to the present problem in that the equation to be estimated, namely $TB = TB(Y_d, Y_f, e)$, is in reduced form, thereby avoiding problems associated with endogenous right-hand-side variables (Pesaran and Shin 1999).

[34] In our model, the import demand elasticity is $|\eta_m| = \frac{|\eta| + k_s\varepsilon_d}{k_m}$, where η = $k_d\eta_d + k_x\eta_x$ is the overall demand elasticity for domestic consumption and exports. For import demand to be price inelastic ($|\eta_m| < 1$), overall demand must be sufficiently price inelastic to satisfy $|\eta| < (k_m - k_s\varepsilon_d)$. This condition would seem difficult satisfy, even in the short-run situation where $\varepsilon_d = 0$. Thus, in instances where LOP and the other assumptions underlying our model hold, one would expect the MLC to hold as well.

The first step is to specify the long-run equilibrium relationship. For this purpose, we adopt the constant elasticity specification

(34) $\ln TB_{i,t} = \alpha + \beta_1 \ln Y_t^{US} + \beta_2 \ln Y_t^{ROW} + \beta_3 \ln ER_{i,t} + \varepsilon_{i,t}$,

where *i* indexes the type of agricultural product (1 = bulk, 2 = high value, and 3 = combined bulk and high value); *t* indexes the year (1976-2012); TB_{it} is the U.S. trade balance defined as the real value of U.S. exports divided by the real value of U.S. imports; Y_t^{US} is real U.S. per capita GDP; Y_t^{ROW} is real per capita GDP for world less US (in U.S. dollars); ER_{it} is the real trade weighted exchange rates between the United States and the currency of foreign trading partners for U.S. bulk products, U.S. high-value products, and combined bulk and high-value products, respectively[35]; and ε_{it} is the error term. The exchange-rate variables are defined as U.S. dollar divided by Foreign Currency Unit (*ER* = DCU/FCU). Hence, an increase in *ER* implies domestic currency *devaluation*.

The betas are long-run elasticities. Theory predicts $\beta_1 < 0$ (a rise in domestic income will increase imports and reduce exports, causing *TB* to decline). Intuitively, a rise in foreign income and a weaker domestic currency should each improve the trade balance, implying positive signs for β_2 and β_3. However, as we saw in the theoretical analysis, intuition is correct only under certain conditions. The signs of β_2 and β_3 are an empirical issue.

The second step is to write the long-run equilibrium relation in ARDL form

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[35] According to the Economic Research Service (2001), nominal exchange rates are those observed and are a result of the market and other forces out of our control. Real exchange rates are nominal rates adjusted for inflation. Trade weighted exchange rates are calculated with a trade-weight index. These indices are constructed by multiplying the average trade weight of a country in U.S. exports, exports to the world, and U.S. imports. These weights are average dollar shares of U.S. exports, exports to the world, and U.S. imports for the relevant commodity. The current exchange rate for each country (in units per dollar) is then adjusted by taking the ratio of the same period CPI in the U.S. to the country in question. The percent change from the base period is then multiplied by the weight. These weighted changes are summed into a total, which is the “real” index.

(35) $\Delta \ln TB_{i,t} = a + \sum_{k=1}^{p} b_k \Delta \ln TB_{i,t-k} + \sum_{k=1}^{p} c_k \Delta \ln Y_{t-k}^{US} + \sum_{k=1}^{p} d_k \Delta \ln Y_{t-k}^{ROW} + \sum_{k=1}^{p} e_k \Delta \ln ER_{i,t-k} + \gamma_1 \ln TB_{i,t-1} + \gamma_2 \ln Y_{t-1}^{US} + \gamma_3 \ln Y_{t-1}^{ROW} + \gamma_4 \ln ER_{i,t-1} + d_1 + d_2 + d_3 + d_4 + u_{i,t}$

where Δ is the difference operator, p is the lag order, *dummy* variable of d_1 represents 1 during the recent recession period of 1980 to 1982 otherwise 0, and d_2 represents 1 within the second recession period of 1990-1991 otherwise 0; d_3 would be 1 during the third recession at the year of 2001 otherwise 0; d_4 indicates 1 in the most recent recession period from 2007 to 2009 and otherwise 0; $u_{i,t}$ is a serially uncorrelated error term. The F-test will be used for examining if there are different effects among the decennial recessions during our sample period (the null hypothesis is:$d_1 = d_2 = d_3 = d_4 = 0$).

Furthermore, the terms involving the gamma parameters constitute the error-correction term EC_{t-1}. Equation (35) is called the error-correction form of the ARDL model (Baek *et al*., p. 218). The gammas represent the long-run (cointegrating) relationship. The coefficients of following the summation symbols indicate short-run effects. The existence of a level relationship (cointegration) is determined by testing the null hypothesis $\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = 0$ using the asymptotic F-values tabulated by Pearson, Shin and Smith (2001). If the null hypothesis is rejected, the variables are cointegrated. If test results are inconclusive, the ARDL model is re-specified with an error correction term. If the estimated coefficient of EC_{t-1} is negative and significant, the variables in the trade balance relation are said to be cointegrated. The tests are valid even if the regressors are not integrated to the same order (e.g., $I(0)$ or $I(1)$). For details, see Baek *et al*. (2009) and the references therein.

The empirical analysis is completed by estimating separate equations for exports and imports. The ARDL specifications are identical to equation (35) except that the *TB* variable is replaced with *EX* and *IM* to represent, respectively, export value and import value.

3.4 The Data and Testing Procedure

We apply annual data over the period 1976 to 2012. Because of the limitations of the data to annual series, the more dynamic aspects that would be present in quarterly or monthly data may not be identified at this stage. The total values of exports and imports for agricultural products between the United States and rest of the world (ROW) are collected from the Foreign Agricultural Trade of the United States (FATUS) database of the U.S. Department of Agriculture (USDA). Different from the previous work, the U.S. trade balance in this analysis is expressed as the ratio of real value of U.S. exports to real value of U.S. imports with the ROW. In that it is not sensitive to the units of measurement by using the ratio and hence can be interpreted as the real trade balance (Baek et al., 2009); meanwhile, the ratio can narrow the range of the variable to make it less susceptible to outlying or extreme observations (Wooldridge 2000).

The real gross domestic product (GDP) index (2005=100) is used as a proxy for the real income of the United States and ROW which is obtained from the Economic Research Service (ERS) and real trade weighted exchange rates for U.S. bulk and high value products are both from FAS while the world U.S. agriculture trade weighted real exchange rate is from ERS International Macroeconomic Data Set. Since the exchange rate is defined as the number of domestic currency (U.S. dollars) per unit of foreign currency in this article, an increase in the exchange rate implies a depreciation of the U.S. dollar from the domestic consumers' perspective and a currency strengthening from the foreign buyers' perspective.

As Pesaran, Shin, and Smith (2001) note, it is crucial to balance between choosing *p* sufficiently large to mitigate the residual serial correlation problems and sufficiently small so that equation (35) is not unduly over-parameterized, particularly in view of the limited time-series data which are available (Pesaran, Shin, and Smith 2001, p. 308). Hence, we employed the Akaike Info Criterion (AIC) for lag selection.

With the selected lag orders, we then test the existence of a level relationship (counteraction) among variables. For this purpose, the null hypothesis of no level relationship, namely (γ_1=γ_2 = γ_3 = γ_4= 0) in equation (35) is tested, irrespective of whether the regressors are purely *I* (0), purely *I* (1), or mutually cointegrated. Pesaran, Shin, and Smith (2001) used an *F*-test with two sets of asymptotic critical values in which all the regressors are assumed to be purely *I* (0) or purely *I* (1). This is called a "bounds testing" procedure since the two sets of critical values provide critical value bounds for all possibilities of the regressors into purely *I* (0), purely *I* (1), or mutually cointegrated (Pesaran, Shin, and Smith 2001, p. 290). With k = 3 for the U.S. bulk, high-value and combined agricultural products, for example, the *F*-statistic value is around 5.0, 4.1, and 10.7 respectively, which all lies outside the upper level of the 10 percent critical bounds (See Table 3.2). As a result, the null hypothesis that there is no cointegrated trade balance equation can be rejected, irrespective of whether the regressors are purely *I* (0), purely *I* (1), or mutually cointegrated. Also, to provide the further evidence of cointegration, the error-correction terms in the ARDL model can be used to determine the existence of cointegrated trade balance equations following Kremers, Ericson, and Dolado (1992) and Banerjee, Dolado, and Mestre (1998). Hence, a negative and significant lagged error-correction term would imply the variables be cointegrated.

3.5 Empirical Results

In economics, the 'J curve' refers to the trend of a country's trade balance following a devaluation or depreciation under a certain set of assumptions[36]. A devalued currency means imports are more expensive, and exports are cheaper but the volume of imports and exports change little immediately, which would cause a depreciation of the current account (a bigger deficit or smaller surplus). Immediately thereafter, the volume of imports and exports may remain largely unchanged due in part to pre-existing trade contracts that have to be honored. Moreover, in the short run, demand for the more expensive imports (and demand for exports, which are cheaper to foreign buyers using foreign currencies) remain price inelastic. This is due to time lags in the consumer's search for acceptable, cheaper alternatives (which might not exist). Thus, the sign of the coefficient of the exchange rate can signal if there exists a J-curve effect or not; and in another words, an initially negative sign followed by a positive one on the lag coefficients would be consistent with the J-curve effect. In table 3.3, none of the coefficient estimates show the negative sign at the current period and positive signs in the following three lag period. Or rather, the last one or two lagged variables show statistical significance for all the three categories, which might imply that in the short run the exchange rate not be the dominating factor in U.S. trade to ROW, and our results do not exactly hold for the J-curve pattern. Likewise, previous studies found that the short-run behavior of the trade balance in response to real exchange rate shocks cannot agree to the J-curve hypothesis either (e.g. Backus et al. 1994, Bahmani-Oskooee and Ratha 2004, Baek et al. 2009); partly it was because the dollar wasn't strong for the reason it normally would have been-because of a strong international trade position. Instead, it was strong because a huge sum of foreign investment money was flowing into the U.S. seeking "safe haven"

[36] See Wiki reference: http://en.wikipedia.org/wiki/J_curve#Balance_of_trade_model.

investments and because of the high rates of return available because of high interest rates in the U.S. compared to the rest of the world. Thus, devaluing the dollar did not change that significantly (Pool and Stamos, 1990). Beyond that, during the period of 1980s, while the dollar did fall against some U.S. trading partners, e.g. Japan and West Germany, it did not fall against others, such as the newly industrialized countries of Taiwan and South Korea and the Pacific Rim countries[37]. Before the first recession within our sample period, the exchange rate line reached peak for all the three groups, which means the dollar against ROW currency was devalued to the most extent, and the U.S. trade surplus of bulk and combined goods hit the highest point in the following year. However, till the last year of the sample period, the trade surplus of combined agricultural products fell by 34 percent compared to its initial value, and the value of bulk goods even decreased about 70 percent though going through ups and downs. Only the trade deficit of U.S. high-value products improved by 36 percent over the past four decades (See Figure 3.1).

It should also be pointed out that the coefficients of the error-correction terms are all negative for bulk products, high-value products and the combined agricultural products, and all are statistically significant at least at the level of 10 percent. For the U.S. bulk and combined agricultural products, the speed of adjustment towards equilibrium would both be around 80 percent, and for high value products, the speed towards equilibrium is about 30 percent, which further provides evidence of the existence of the long-run relationship among variables (Kremers, Ericson, and Dolado 1992, Banerjee, Dolado and Mestre 1998, and Baek et al.2009). The

[37] Since 2000, U.S. goods imports from developing countries have exhibited higher growth (almost 4 times as much) than that from industrial countries, 130 percent compared with 36 percent. On the other hand, U.S. goods exports to developing countries have grown almost three times as fast as U.S. goods exports to industrial countries, 135 percent compared to 54 percent. Due to this long-term higher-growth difference, the share of U.S. goods exports to developing countries have grown from 45 percent in 2000 to 55 percent in 2011. For example, the percentage change of U.S. exports to China reached 554% versus that of U.S. imports from China was 300.2% from 2000-11.

findings further justify the ARDL modeling of U.S. agricultural trade of all the three categories with ROW, in which the F-statistics results are cointegration (See Table 3.2).

Table 3.4 yields the estimated results from the full model before any F-tests for recession effects, and the coefficient estimates of trade weighted real exchange rate carry the positive relationship with the U.S. trade balance for all the three products; at meanwhile, high-value and combined agricultural products are statistically significant at the 5 percent level (Table 3.4). It states that the depreciating domestic currency (U.S. dollars) lead to an improvement of U.S. trade balance in the long term, meaning with each one percent of U.S. dollar depreciation, U.S. trade balance of high-value commodities and total agricultural products would enhance by 1.62 percent and 1.88 percent respectively. Besides, for both of high value and combined products, the U.S. trade balance displays the positive relationship with domestic income and negative with foreign income, where only trade balance of combined goods with foreign income shows statistical significance at the level of 5 percent. According to Baek et al. (2009), it demonstrates that a rise of real domestic income increases the domestic demand for U.S. exports while decreases the demand for foreign imports, thereby improving the U.S. trade balance; and vice versa, for the foreign income, it has adverse relationship with the U.S. trade balance and a deteriorating effect accordingly. One of the possible explanations for the finding is that, since imports are defined as the difference between domestic consumption and production, an increase in domestic income could increase the domestic production of import-substitute commodities faster than a rise in domestic consumption, thereby leading to the reduction of domestic imports[38]; meanwhile, records show HVP exports exceeded exports of bulk products at the beginning of year 1990 for the first time and kept the trend until today, then the income effects of

[38] See Magee (1973), Bahmani-Oskooee (1985), and Bahmani-Oskooee and Ratha (2004)

HVP could be the major contributor and thus one percent increase of foreign GDP, *ceteris paribus*, would decrease the U.S. trade surplus of combined agricultural goods by around 1.32 percent[39]. Furthermore, our results found that growth in incomes has a larger effect on exports of bulk commodities than on high-value exports, which is opposite to the finding drawn by Shane et al. (2009).

Compared to the recent work by Baek and Koo (2011), our results indicate the similar results on the U.S. import side but different on the export side in the long run. That is, the U.S. exports are very responsive to exchange rate changes for bulk, high-value (consumer-oriented products) and the combined commodities but do not seem to respond much to income; And for U.S. imports, both of bulk and high-value products are relatively insensitive to exchange rate changes.

As for dummy variables of recessions, F-test indicates that only trade balance model and import model of U.S. combined agricultural products reject the null hypothesis that the four dummy variable coefficients are zero, which means the U.S. recessions occurred in the past forty years have had different effects on trade balance as well as imports of U.S. combined products. And the dummy variable effects imply that the U.S. trade balance would decrease by 0.21 percent during the 1980s crisis compared with other years, while the U.S. imports would increase by 0.15 percent within the recent great recession from 2008 to 2009. This might suggest that when the first economic recession of our sample period occurred, the U.S. trade balance of combined agricultural goods decreased, meaning U.S. exports declined or U.S. imports increased

[39] Stephen (ERS, USDA, 2013) introduced the market share of middle-income and high-income countries that import from U.S. Agricultural products, and they now account for the largest share of U.S. agricultural exports of both bulk products and semi-processed high-value products (*e*.g. wheat, soybeans, and soybean meal); while for the other categories of high-value products---raw products and processed products, high-income countries remain the largest U.S. markets, followed by the upper middle-income countries.

or both. One possible reason should be correlated with decreasing U.S. GDP declining: Table 3.4 demonstrates the U.S. income is positively related with its trade balance and negatively related with its imports for combined products, therefore U.S. trade balance would be hurt and U.S. imports would increase under such a circumstance. Another reason could be the U.S. dollar was getting strong instead of weakening (or devaluation) in both recession periods, the 1980s recession and the most recent recession, during which the exchange rate curve of combined agricultural goods started to slump seen from Figure 3.2, thus, the U.S. exports became decreasing and imports increasing correspondingly. Besides, the most recent recession was verified to positively impact on U.S. imports, which further states that the recession could increase the imports. And our results agree to the previous work by Liefert and Shane (2009) mentioned earlier. For individual agricultural goods such as bulk or high-value products, F-tests indicate that none of the recessions has any significant impact on U.S. trade balance. This might imply that global trade partners have tended to constrain exports/imports to the U.S. combined agricultural products rather than to individual ones in response to recessions, and basically the U.S. recession effects for agricultural trade in the last forty years were less than expected and the U.S. monetary policy or financial tools for mitigating the recession downturns should have worked to certain extent.

Specifically, the export and import equations are estimated separately for further analysis. The results suggest that the U.S. exports have a positive relationship with the exchange rate in long-run for three cases, and all of the estimates are statistically significant at least at level of 5 percent, highly suggesting that the devaluing domestic currency (U.S. dollars) would increase the domestic (U.S.) exports; meanwhile, in the long term, the real exchange rate is passed through to the U.S. exports of combined agricultural products to the largest degree, compared with individual bulk and high-value goods. Besides, for bulk commodities the relationship between U.S. exports and foreign income appears positive and statistically significant at the level of 10 percent, rather than the other two cases; and one percent increase of foreign per capita GDP would enhance U.S. bulk exports by around 3.57 percent, while one percent of increase of traded weighted real exchange rate would improve U.S. exports by 1.19 percent, which corroborates the conclusion drawn by Shane et al. (2009) that the net effect on total agricultural exports depends on the magnitude of growth in income compared to the magnitude of change in the trade-weighted exchange rate.

By contrast, the reduced import model implies that the real exchange rates of all the three categories are negatively related with the U.S. imports, but only the coefficient of combined agricultural goods shows significance at 10 percent level, corroborating the common-held belief that the devalued U.S. dollar would cause U.S. imports to decline, which is consistent with our hypothetical expectations. The decomposed export model and import model demonstrate the depreciation effects of exchange rate on U.S. trade outweigh the inconclusive effects of either domestic or foreign income in the long run.

As beginning with the pioneering studies of Bahmani-Oskooee and Ardalani (2006, 2007), there has been a growing body of literature that argues that trade balance study could suffer from aggregation bias of data due to the fact that a country tends to export and import different commodities to/from different trading partners, thus in this article we did the *Wald* tests to determine whether or not bulk and high-value commodities can be aggregated across three models (See table 3.5). The results show that all the three null hypotheses are rejected at 5% probability level, which means there might not exist any aggregation bias between bulk and high-value in our study.

As for the F-tests for determining whether recession effects are jointly significant, only combined agricultural goods receive some support in that two of the nine tests show a significant relationship between U.S. import value or trade balance and the recession effects during the past decades. However, the remaining tests show no significance (See Table 3.6).

Though not all the estimated coefficients for the three cases are statistically significant, the signs between exchange rates and U.S. trade balance are the same; according to the earlier mentioned theoretical deductions, the relationship between foreign income and U.S. trade balance is ambiguous and cannot be inferred. Thus, our results might provide empirical evidence, if any, for the previous theoretical argument. Generally speaking, the estimates of real exchange rates present positive relationship with U.S. trade balance not only for individual agricultural commodities of bulk and high-value goods but also for the combined agricultural products within the sample period since late 1970s.

Finally, to test if the estimated coefficients are stable or not over time, we use the cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests to the residuals of ECMs (Eqs. (35). For stability of all estimated coefficients, the plot of these two statistics should stay within the 5% significance level. The overall results of stability test suggest that the estimated coefficients of all models are generally stable over the sample period[40].

3.6 Conclusions

The article applies the dynamic ARDL model of error correction version, not only investigating if there is J-curve effect in the short-run or not, but also taking a deep analysis for U.S. recession effects and exchange rate as well as income growth effects in the long run on the U.S. trade

[40] According to Baek and Koo (2011), these tests are known to have low power and could miss important breaks.

balance of agricultural commodities which mainly consists of bulk products and high-value products. Our results indicate that there is no significant J-curve effect for three cases, while the long-run effect demonstrates that the domestic currency devaluation is positively related with U.S. agricultural trade balance for bulk, high-value and combined agricultural products, though the high-value products appear the more modest effects compared to the other two.

In sum, the real trade-weighted exchange rate is found to be the key determinant of U.S. agricultural trade balance in the long-term, rather than domestic or foreign income. We find that the three categories of agricultural products do indeed respond differently to exchange rate and income. For bulk and high-value products, U.S. exports are highly sensitive to exchange rate and foreign income, while U.S. imports barely respond. For combined agricultural products, on the other hand, U.S. exports respond greatly to exchange rate, and U.S. imports behave significantly with respect to both of changes in exchange rate and foreign income; besides, the 1980s recession had significant effects on U.S. trade balance while the most recent recession had great impact on U.S. imports, showing the U.S. trade with ROW partners was mainly influenced by the two times economic crisis during our sample period.

To our knowledge, it is the first time that the ARDL model is applied for the recession effects on individual groups of U.S. total agricultural commodities: Bulk and High-Value products. Different from the previous work by Baek and Koo (2011), this paper analyzes both of dollar devaluation impacts and the decennial U.S. economic recession effects on U.S. trade balance by employing the aggregated data and explores the reasons that the J-curve effect did not show up during the past decades; also, the test shows that there is no any aggregation bias between bulk and high-value in our study, making both of the dynamic short-run effect and stable long-run effect estimated from the reduced model at linkage with the theoretical structure equations more solid and convincing.

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Appendix 1: Tables and Figures for Chapter I

Table 1.1 Descriptive statistics of monthly data, Oct. 2005 to Dec. 2011

Variable	Mean	Std. Dev.	Min	Max
Liquid milk consumption (kg/capita)	1.734	0.211	1.340	2.112
Powdered milk consumption (kg/capita)	0.044	0.006	0.031	0.062
Yogurt consumption (kg/capita)	0.410	0.075	0.246	0.620
Retail liquid milk price (yuan/kg)[a]	6.903	1.462	4.676	9.745
Retail powdered milk price (yuan/kg)	94.475	31.078	48.136	189.550
Retail yogurt price (yuan/kg)	8.193	1.367	5.726	11.261
Dairy expenditure (yuan /capita)	19.216	3.202	13.536	25.686
Liquid milk expenditure share	0.558	0.047	0.470	0.645
Powdered milk expenditure share	0.192	0.047	0.120	0.282
Yogurt expenditure share	0.158	0.024	0.093	0.194
Media coverage index	4.381	8.758	0.000	39.600

a. The money unit is RMB yuan.

Table 1.2 Coefficient Estimates for Three Dairy Products in China

Explanatory variables	Liquid milk	Powdered milk	Yogurt
Price			
Liquid milk	–1.06** [a]	0.07**	
	(0.08) [b]	(0.007)	
Powdered milk	0.39**	–0.02**	
	(0.05)	(0.004)	
Yogurt			–0.05**
			(0.02)
Total expenditure	1.78**	0.02	0.63**
	(0.12)	(0.01)	(0.07)
Media Coverage			
DV	–0.03**	–0.0001	0.01
	(0.01)	(0.001)	(0.008)
N_t	–5.22e–05	–6.31e–05	–0.0008**
	(0.000446)	(4.16e–05)	(0.000257)
N_{t-1}	0.0007	–1.67e–05	0.0002
	(0.000514)	(4.79e–05)	(0.000296)
N_{t-2}	–1.77e–05	–3.66e–06**	1.50e–05*
	(1.26e–05)	(1.17e–06)	(7.24e–06)
N_{t-3}	6.09e–09	–1.40e–07**	6.79e–07**
	(3.27e–07)	(3.05e–08)	(1.88e–07)
Seasonality			
Jan	–0.05**	0.003*	–0.05**
	(0.01)	(0.001)	(0.009)

Feb	–0.004	0.0008	0.0002
	(0.02)	(0.001)	(0.009)
Mar	–0.03*	0.0001	0.03**
	(0.02)	(0.002)	(0.01)
Apr	–0.21**	0.0003	0.04**
	(0.02)	(0.002)	(0.01)
May	–0.20**	–0.002	0.06**
	(0.02)	(0.002)	(0.01)
Jun	–0.17**	–0.005**	0.087**
	(0.02)	(0.002)	(0.01)
Jul	–0.34**	–0.006**	0.13**
	(0.02)	(0.002)	(0.01)
Aug	–0.28**	–0.006**	0.09**
	(0.02)	(0.002)	(0.01)
Sep	–0.27**	–0.005*	0.04**
	(0.02)	(0.002)	(0.01)
Oct	–0.002	–0.004**	0.05**
	(0.02)	(0.001)	(0.009)
Nov	–0.02	–0.001	0.01
	(0.01)	(0.001)	(0.008)
TREND	–0.004**	–0.0002**	0.0007
	(0.0008)	(7.04e–05)	(0.0004)
Intercept	–1.11**	–0.03	–0.38*
	(0.32)	(0.03)	(0.18)
R-squared	0.99	0.88	0.97
D–W	1.59	1.20	1.32

[a.] Single asterisk indicates significance at 95% confidence level; double asterisk indicates significance at 99% confidence level. [b.] Standard errors are in parentheses.

Table 1.3 Estimated Price and Expenditure Elasticites

	– – – – – – – – – – – – Price elasticity[a] – – – – – – – – –			Expenditure elasticity [a]
	Liquid milk	Powdered milk	Yogurt	
Liquid milk	–0.63	1.71		1.08
Powdered milk	0.23	–0.43		0.50
Yogurt			–0.15	1.71

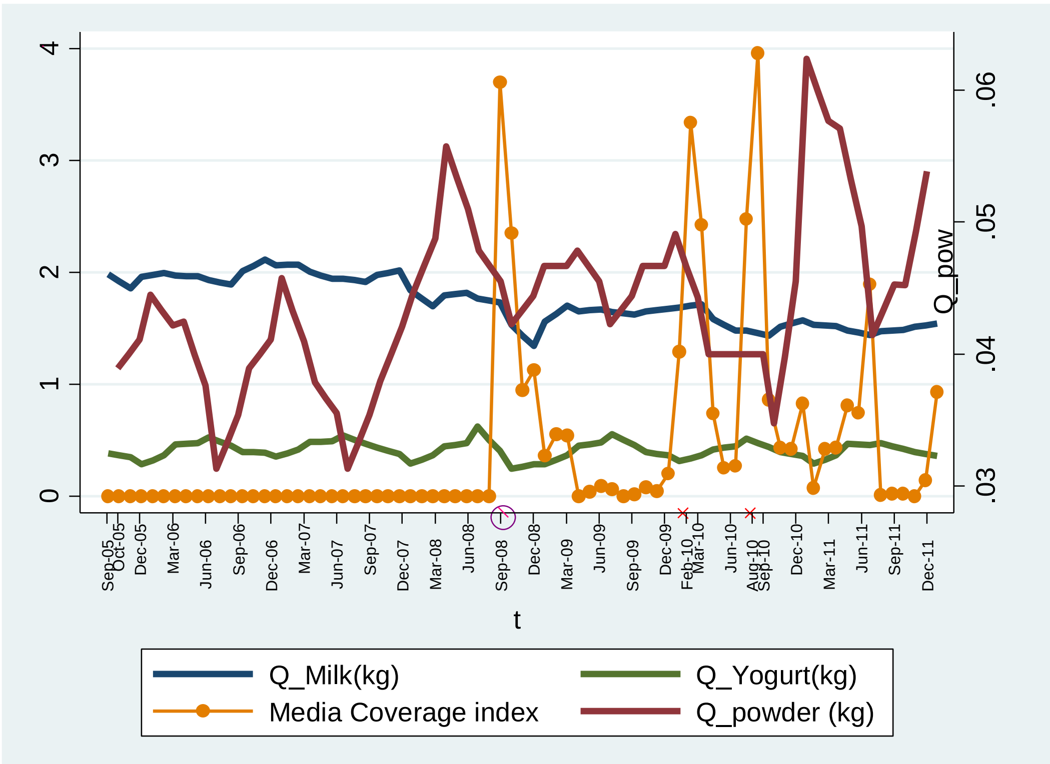
[a.] All elasticities are evaluated at sample means.

Table 1.4 Simulated Annual Per Capita Dairy Sales without Media Coverage on the Melamine Incident, China

	Liquid milk demand			Powdered milk demand			Yogurt demand		
Year	Fitted (a)	Simulated (b)#	Δ (a)–(b)	Fitted (a)	Simulated (b)	Δ (a)–(b)	Fitted (a)	Simulated (b)	Δ (a)–(b)
2009	19.76	20.17	–0.41	0.55	0.57	–0.020	5.00	4.75	0.25
2010	18.68	18.69	–0.01	0.5	0.57	–0.070	4.85	4.63	0.22
2011	17.97	18.36	–0.39	0.62	0.61	0.010	4.84	4.82	0.02
Difference				kg/capita					
Δ(2010–2009)	–1.08	–1.48	0.40	–0.05	0.00	–0.05	–0.15	–0.12	–0.03
Δ(2011–2009)	–1.79	–1.81	0.02	0.07	0.04	0.03	–0.16	0.07	–0.23
				Percent Change					
Δ(2010–2009)	–5.47	–7.34	1.87	–9.09	0.00	–9.09	–3.00	–2.53	–0.47
Δ(2011–2009)	–9.06	–8.97	–0.08	12.73	7.02	5.71	–3.20	1.47	–4.67

Notes: # Fitted values with the zero coverage media index in 2009 for dairy products.

Figure 1.1 Monthly per Capita Dairy Demand and Media Coverage Indices, China 2005–11



Notes:

The powdered milk demand is shown on the right *y*-axis.

The original media coverage indices are all divided by 10 shown on the graph.

O The official media started to report the melamine incident nationwide in Sept. 2008.

× The original media coverage index is above 30.

Appendix 2: Tables and Figures for Chapter II

Table 2.1 Example of Segmented Variable Using the First Twelve Observations of the Whole Milk Farm Price Variable

Date	**Farm Price (P_{ft})**	ΔP^{+}_{ft}	ΔP^{-}_{ft}	FR_t	FF_t
Jan--96	1.2768	0	0	0	0
Feb	1.2800	0.0032	0	0.0032	0
Mar	1.2664	0	–0.0136	0.0032	–0.0136
Apr	1.2568	0	–0.0096	0.0032	–0.0232
May	1.2680	0.0112	0	0.0144	–0.0232
Jun	1.3248	0.0568	0	0.0712	–0.0232
Jul	1.3712	0.0464	0	0.1176	–0.0232
Aug	1.3832	0.0120	0	0.1296	–0.0232
Sep	1.4312	0.0480	0	0.1776	–0.0232
Oct	1.4704	0.0392	0	0.2168	–0.0232
Nov	1.5040	0.0336	0	0.2504	–0.0232
Dec	1.4128	0	–1.1400	0.2504	–1.1632

Note: $\Delta P^{+}_{ft} = P_{ft} - P_{ft-1}$ if $P_{ft} > P_{ft-1}$

$= 0$ otherwise

$\Delta P^{-}_{ft} = P_{ft} - P_{ft-1}$ if $P_{ft} < P_{ft-1}$

$= 0$ otherwise

FR_t= cumulative sum of ΔP^{+}_{ft}

FF_t=cumulative sum of ΔP^{-}_{ft}

Table 2.2 Statistical Summary of Whole Milk, United States Data, January 1996-December 2011.

Variable	Retail Units	N	Mean	Std. Dev.	Min	Max
Retail Price	$/gal.	192	3.0393	0.3722	2.5320	3.9610
Farm Price	$/gal.	192	1.4530	0.2676	1.0968	2.1584
Rising	$/gal.	192	0.0359	0.0597	0	0.48
Falling	$/gal.	192	–0.0500	0.1926	–2.25	0
Market Cost	Index	192	540.5969	64.9547	448.3	656

Table 2.3 Granger Causality Test from 1996:01 to 2011:12 based on Monthly Data of Farm and Retail prices for national Whole Milk

Effect	Hypothesized	F-	p-value
Farm price	Retail price	2.20	0.1138
Retail price	Farm price	5.18	0.0065*

Note: The null hypothesis is that one series does not Granger cause another. The Granger causality tests use a lag length of two months. *indicates statistical significance at the 0.05 level. EVIEWS 7.0 was the statistical package employed to conduct the Granger tests.

Table 2.4 Empirical Results of the Johansen Cointegration Tests for Farm prices and Retail Prices of Whole Milk.

Hypothesized No. of cointegrated equations	Trace test Statistic	p-value	Maximal test Statistic	p-value
None *	19.82	0.0104*	17.30	0.0161*
At most 1	2.53	0.1118	2.53	0.1118

Note: The intercept (no trend) option with three lags was used in conjunction with these tests. The level of significance chosen for this analysis was 0.05. *indicates statistical significance. EVIEWS 7.0 was the statistical package employed to conduct the cointegration tests.

Table 2.5 Empirical Results of the Houck Procedure for U.S. Whole Milk Based on 1996-2011 Monthly Data

	Model I	Model II	Model III	Model IV
Trend	–0.0006		–0.0003	
	(0.0004)		(5.73E-05)	
ΔP_{ft}^{+} (SR^{+}) [a]	0.9099 [b]	0.8982	0.9056	0.8472
	(0.0582) [c]	(0.0580)	(0.0580)	(0.0586)
ΔP_{ft-1}^{+}	0.2272	0.2129	0.2217	0.1571
	(0.0433)	(0.0427)	(0.0429)	(0.0426)
ΔP_{ft-2}^{+}	–0.0762	–0.0865	–0.0802	–0.1253
	(0.0425)	(0.0421)	(0.0422)	(0.0418)
LR+	1.0609	1.0245	1.0471	0.8790
	(0.1038)	(0.1023)	(0.1028)	(0.1027)
ΔP_{ft}^{-} (SR^{-}) [a]	0.0568	0.0583	0.0574	0.0647
	(0.0180)	(0.0181)	(0.0180)	(0.0186)
ΔP_{ft-1}^{-}	0.0279	0.0299	0.0288	0.0362
	(0.0126)	(0.0127)	(0.0126)	(0.0136)
ΔP_{ft-2}^{-}	0.0089	0.0105	0.0096	0.0146
	(0.0129)	(0.0129)	(0.0129)	(0.0134)
LR–	0.0936	0.0987	0.0958	0.1155
	(0.0298)	(0.0300)	(0.0298)	(0.0328)
MD_t [a]	0.0003	–0.0002		
	(0.0003)	(5.75E-05)		
AR(1)	0.2329	0.2445	0.2361	0.3264
	(0.0757)	(0.0755)	(0.0755)	(0.0753)
AR(2)	0.0761	0.0855	0.0788	0.1452
	(0.0754)	(0.0751)	(0.0751)	(0.0757)
R square	0.6434	0.6380	0.6417	0.6040
D.W.	1.99	1.99	1.99	2.01
AIC	–3.1680	–3.1637	–3.1741	–3.0848
SIC	–3.0277	–3.0410	–3.0514	–2.9796

Note: [a] ΔP_{ft}^{+} and ΔP_{ft}^{-} are *identified above*, where *ft-i* and *rt-i* implies lag lengths (*i=0, 1, 2, 3*); where MD_t is marketing cost variables, expressed as deviations from its initial value. *The dependent value is expressed as the first difference of retail price.*

[b] *Parameter estimate*

[c] *Standard errors in parentheses*

Table 2.6 Error Correction Model Results for U.S. Whole Milk Based on 1996-2011 Monthly Data

	Model I	Model II	Model III	Model IV
Trend	–			
0.0005			–2.59E-05	
	(0.0002)		(4.95E-05)	
ΔP_{ft}^{+} (SR^{+}) [a]	0.9316[b]	0.9191	0.9250	0.9178
	(0.0573) [c]	(0.0576)	(0.0577)	(0.0559)
ΔP_{ft-1}^{+}	–0.0151	–0.0391	–0.0275	–0.0418
	(0.0889)	(0.0891)	(0.0895)	(0.0851)
LR+	0.9165	0.8800	0.8975	0.8760
	(0.0944)	(0.0938)	(0.0948)	(0.0853)
ΔP_{ft}^{-} (SR^{-}) [a]	0.0469	0.0453	0.0453	0.0454
	(0.0167)	(0.0169)	(0.0169)	(0.0169)
ΔP_{ft-1}^{-}	0.0176	0.0183	0.0182	0.0183
	(0.0103)	(0.0104)	(0.0104)	(0.0104)
ΔP_{ft-2}^{-}	0.0020	0.0031	0.0031	0.0031
	(0.0116)	(0.0117)	(0.0117)	(0.0117)
LR–	0.0665	0.0667	0.0666	0.0668
	(0.0237)	(0.0240)	(0.0239)	(0.0239)
ECT_{t-1}^{+}	–0.2467	–0.2476	–0.2377	–0.2501
	(0.0400)	(0.0404)	(0.0401)	(0.0323)
ECT_{t-1}^{-}	0.0004	0.0263	0.0248	0.0264
	(0.0338)	(0.0318)	(0.0320)	(0.0318)
ΔP_{rt-1}	0.1087	0.1266	0.1237	0.1271
	(0.0722)	(0.0725)	(0.0725)	(0.0721)
ΔP_{rt-2}	–0.0551	–0.0542	–0.0519	–0.0548
	(0.0481)	(0.0486)	(0.0485)	(0.0481)
ΔP_{rt-3}	–0.1370	–0.1373	–0.1335	–0.1383
	(0.0459)	(0.0463)	(0.0463)	(0.0453)
MD_t [a]	0.0005	–5.09E-06		
	(0.0002)	(4.86E-05)		
R square	0.7117	0.7043	0.7048	0.7043
D.W.	1.98	2.00	1.99	2.00
AIC	–3.3343	–3.3198	–3.3212	–3.3304
SIC	–3.1443	–3.1470	–3.1485	–3.1749

Note: [a] ΔPft^{+} and ΔPft^{-} are identified above, where f_{t-i} and r_{t-i} implies lag lengths (i=0, 1, 2, 3); where MD_t is marketing cost variables, expressed as deviations from its initial value. The dependent value is expressed as the first difference of retail price.

[b] Parameter estimate

[c] Standard errors in parentheses

Table 2.7 Elasticites of Farm-Retail Price Transmission for Whole Milk under Rising and Falling Farm Prices, United States, Based on Data Covering January 1996-December 2011

	Elasticity When Farm Prices Are:			
	Rising		Falling	
	LR	SR	LR	SR
Houck approach	0.5006	0.4422	0.0458	0.0217
ECM approach	0.4291	0.4329	0.0318	0.0274

Note: The elasticites are evaluated on the mean value.

Table 2.8 Theoretical Values of Price Transmission Elasticities under Retail Demand versus Farm Supply Shifts (Gardner 1975, p.403, table 1)

Parameters					Transmission Elasticity When the Margin Change is due Exclusively to a:	
σ	e_a	e_b	η	S_b	Retail Demand Shift	Farm Supply Shift
0.5	1.0	2.0	–0.5	0.5	0.80	0.50
0	1.0	2.0	–0.5	0.5	0.75	1.44
0	1.5	2.0	–0.5	0.5	0.88	0.44
0	2.0	2.0	–0.5	0.5	1.00	0.44
0	2.0	1.0	–0.5	0.5	1.50	0.40
0	1.0	2.0	–1.0	0.5	0.75	0.40

Table 2.9 Hausman Test for Exogeneity of Food Marketing Costs

Equation	*t*-statistic	*p*-values	Result
Whole Milk *absolute* marketing margin	1.16	0.9483	Fail to reject
Whole Milk *relative* marketing margin	0.86	0.9731	Fail to reject
Whole Milk price transmission	0.86	0.9731	Fail to reject

Table 2.10 Estimated Marketing Margin and Price Transmission Relations for Whole Milk Based on 1996-2011 Quarterly Data

VARIABLES/ Statistic	***Absolute* Marketing Margin**		***Relative* Marketing Margin**	
	(i)	(ii)	(iii)	(iv)
Farm Price	–0.050		–0.557***	
	(0.040)		(0.040)	
Farm Price*Dummy	–0.145**		–0.167**	
	(0.068)		(0.068)	
Marketing Cost	0.573*	0.330	–0.457	–0.457
	(0.314)	(0.388)	(1.218)	(1.218)
Consumer Income	0.309	0.373	0.577	0.577
	(0.269)	(0.335)	(1.053)	(1.053)
Corn Price	0.006	0.008	0.014	0.014
	(0.016)	(0.020)	(0.063)	(0.063)
Constant	–0.008*	–0.002	–0.001	–0.001
	(0.004)	(0.004)	(0.012)	(0.012)
Observations	63	63	63	63
R-squared	0.409	0.047	0.938	0.009
D.W.	1.85	2.10	1.80	2.22

Note: Standard errors in parentheses (*** p<0.01, ** p<0.05, * p<0.1);
Dummy=1 while Farm price is falling, otherwise equals 0.

VARIABLES/ Statistic	Price Transmission	
	(v)	(vi)
Farm Price	0.443***	0.441***
	(0.040)	(0.041)
Farm Price*Dummy	–0.167**	–0.155**
	(0.068)	(0.069)
Marketing Cost	0.535*	
	(0.313)	
Income	0.265	0.307
	(0.268)	(0.271)
Corn Price	0.008	0.012
	(0.016)	(0.016)
Constant	–0.008*	–0.005
	(0.004)	(0.004)
Observations	63	63
R-squared	0.835	0.827
D.W.	1.80	1.68

Note: Standard errors in parentheses (*** $p<0.01$, ** $p<0.05$, * $p<0.1$)
Dummy=1 while Farm price is falling, otherwise equals 0.

Table 2.11 Tests for Competition followed by Lloyd et al. (2009)

Product	H_0:D=S=0	H_0:D=0	H_0:S=0
Whole Milk	0.80	0.98	0.26
	(0.452)	(0.326)	(0.611)

Figures in brackets are asymptotic *p*-values;

Table 2.12 Wald Tests for Competitive Market Clearing in the U.S. Whole Milk Marketing Channels

Test[a]	Chi Square	*p*-values	Result
Weak-Form	0.83	0.6596	Fail to reject
Strong-Form	4.15	0.3855	Fail to reject

[a] The weak-form test imposes the restriction that the ***absolute*** marketing margin be invariant to shifts in retail demand and farm supply. The strong-form imposes the additional restrictions that the ***absolute*** marketing margin be invariant to farm price, and the elasticity of farm-retail price transmission be equal to the farmers' share of the food dollar. The farmers' share of the consumer dollar for whole milk was set to 0.48 according to USDA estimates (Jones, Dunham).

Table 2.13 Wald Tests for Competitive Market Clearing in the U.S. Whole Milk Marketing Channels

Test[a]	Chi Square	*p*-values	Result
Weak-Form	0.05	0.9759	Fail to reject
Strong-Form	6880.97	0.0000	Reject

[a] The weak-form test imposes the restriction that the ***relative*** marketing margin be invariant to shifts in retail demand and farm supply. The strong-form imposes the additional restrictions that the ***relative*** marketing margin be invariant to farm price, and the elasticity of farm-retail price transmission be equal to the farmers' share of the food dollar. The farmers' share of the consumer dollar for whole milk was set to 0.48 according to USDA estimates (Jones, Dunham).

Figure 2.1 Whole milk, farm and retail prices (U.S. city average)

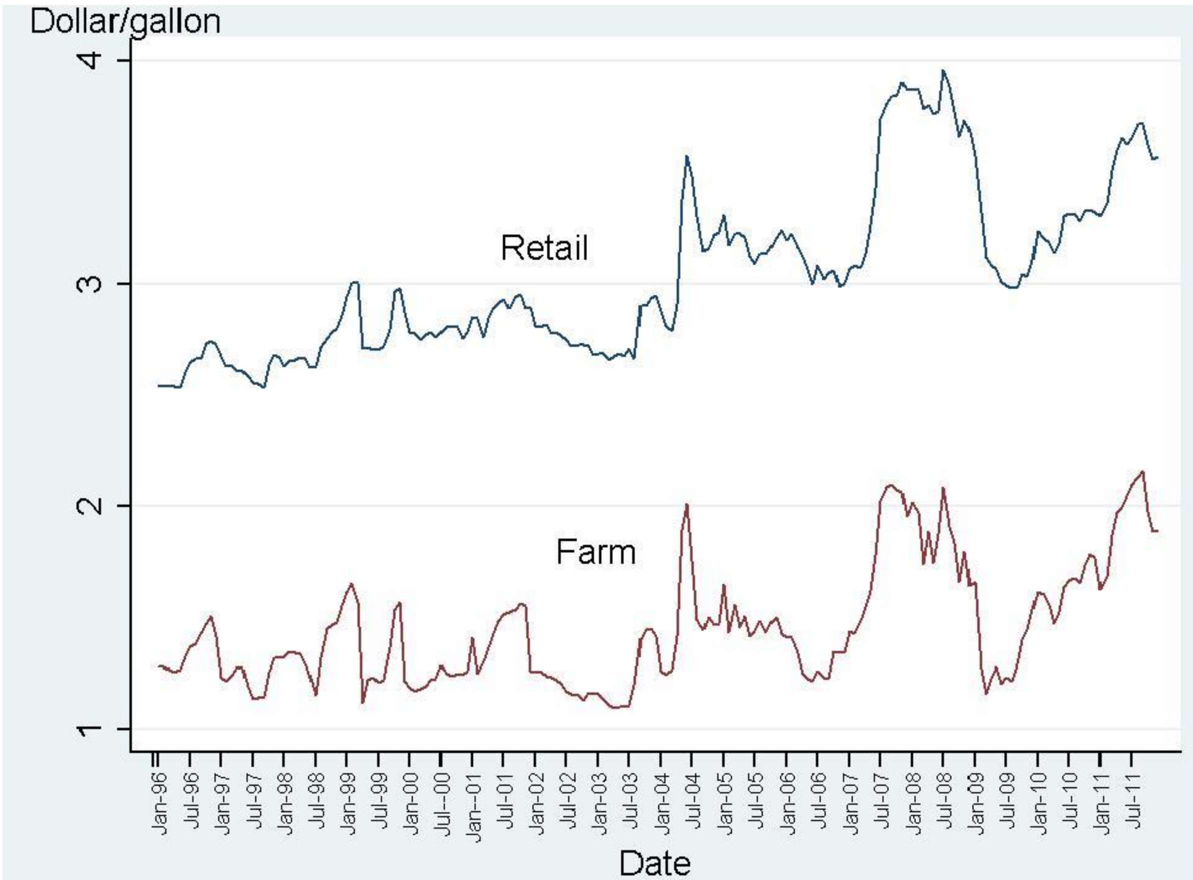


Figure 2.2 The Whole Milk Retail-Producer Price Spread (gallon per dollar)

Note: The graphs are generated by statistical package Stata 12.0.

Appendix 3: Tables and Figures for Chapter III

Table 3.1 The Descriptive Summary of Economic Recession from 1976 to 2012.

Name	Dates	Duration (months)	Peak Unemployment	GDP decline (peak t trough)	Characteristics
Early 1980s recession	1980-1982	1 year 10 months	10.80%	−2.7%	The recession began as the Federal Reserve, under Paul Volcker, raised interest rates dramatically to fight the inflation of the 1970s.The Iranian Revolution sharply increased the price of oil around the world in 1979, causing the 1979 energy crisis. This was caused by the new regime in power in Iran, which exported oil at inconsistent intervals and at a lower volume, forcing prices up. Tight monetary policy in the United States to control inflation led to another recession. The changes were made largely because of inflation carried over from the previous decade because of the 1973 oil crisis and the 1979 energy crisis The early '80s are sometimes referred to as a "double-dip" or "W-shaped" recession.
Early 1990s recession	July 1990-Mar1991	8 months	7.80%	−1.4%	After the lengthy peacetime expansion of the 1980s, inflation began to increase and the Federal Reserve responded by raising interest rates from 1986 to 1989. This weakened but did not stop growth, but some combination of the subsequent 1990 oil price shock, the debt accumulation of the 1980s, and growing consumer pessimism combined with the weakened economy to

					produce a brief recession
Early 2000s recession	March 2001-Nov2001	8 months	6.30%	−0.3%	The 1990s were the longest period of growth in American history. The collapse of the speculative dot-com bubble, a fall in business outlays and investments, and the September 11th attacks brought the decade of growth to an end. Despite these major shocks, the recession was brief and shallow. Without the September 11th attacks, the economy might have avoided recession altogether.
Great recession	Dec 2007-June 2009	1 year 6 months	10.00%	−4.3%	The subprime mortgage crisis led to the collapse of the United States housing bubble. Falling housing-related assets contributed to a global financial crisis, even as oil and food prices soared. The crisis led to the failure or collapse of many of the United States' largest financial institutions: Bear Stearns, Fannie Mae, Freddie Mac, Lehman Brothers, Citi Bank and AIG, as well as a crisis in the automobile industry. The government responded with an unprecedented $700 billion bank bailout and $787 billion fiscal stimulus package. The National Bureau of Economic Research declared the end of this recession over a year after the end date. The Dow Jones Industrial Average (Dow) finally reached its lowest point on March 9, 2009.

See wiki link of
http://en.wikipedia.org/wiki/List_of_recessions_in_the_United_States#Free_Banking_Era_to_the_Great_Depression

Table 3.2 Results of F-Test for Cointegration among Variables of Reduced Trade Balance Model.

Variable	AIC Lags	*F*-statistic	Decision
Bulk	5	5.03	Cointegration
High-Value	4	4.11	Cointegration
Combined	3	10.66	Cointegration

Note: A lag order is chosen based on Akaike Info Criterion (AIC). *F*-statistic for 10 percent critical value bounds is (2.72, 3.77), which is taken from Table CI in Pesaran et al. (2001).

Table 3.3 Coefficient Estimates of Exchange Rate and Error-Correction Terms of the Reduced Trade Balance Model.

Products	Lag Order of Exchange Rate						
	0	1	2	3	4	5	EC_{t-1}
Bulk	−0.56	−0.58	−0.63	−1.60	−1.47**	−1.57*	−0.81**
	(−0.64)	(-0.67)	(−0.78)	(−0.77)	(−2.22)	(−1.95)	(−3.15)
High-Value	0.37	0.06	0.60	−0.57	0.98**		−0.29*
	(1.00)	(0.14)	(1.37)	(−1.45)	(2.75)		(−2.05)
Combined	0.25	−0.24	−0.40	0.69			−0.80**
	(0.53)	(−0.59)	(−1.27)	(1.86)			(−3.91)

Note: ** and * denote significance at the 5 percent and 10 percent levels, respectively. Parentheses are *t*-statistics. *ECt*–1 refers to the error correction term.

Table 3.4 Estimated Reduced-Form Equations for U.S. Agricultural Exports, Imports, and Trade Balance, Annual Data, 1976-2012

	Export Value				Import Value				Trade Balance (Export Value/Import Value)		
	Bulk	High Value	Bulk & High Value		Bulk	High Value	Bulk & High Value		Bulk	High Value	Bulk & High Value
Exchange Rate	1.19** (3.72)[a]	0.73** (3.40)	1.72** (3.60)		−0.00 (−0.00)	−0.39 (−1.34)	−0.54* (−2.09)		1.96 (1.28)	1.62** (2.88)	1.88** (4.96)
U.S. Income	−1.41 (−1.70)	−0.26 (−0.71)	0.23 (0.49)		0.07 (0.04)	−0.22 (−0.50)	−0.60 (−1.71)		−0.07 (−0.03)	0.27 (0.28)	0.18 (0.48)
Foreign Income	3.57* (2.48)	0.16 (0.31)	1.38 (1.38)		0.32 (0.18)	2.36 (1.53)	3.63** (3.16)		0.21 (0.08)	−0.40 (−0.31)	−1.32** (−2.57)
D1	0.01 (0.11)	−0.03 (−0.54)	−0.15 (−1.58)		−0.55 (−1.38)	0.05 (0.71)	0.11 (1.61)		0.77 (1.53)	−0.18* (−1.85)	−0.21** (−2.56)
D2	0.03 (0.32)	−0.00 (−0.03)	−0.03 (−0.49)		−0.46* (−2.09)	−0.04 (−0.74)	−0.01 (−0.36)		0.43* (1.84)	0.04 (0.50)	0.08 (1.37)
D3	0.10 (0.85)	0.12* (2.11)	0.12 (1.49)		0.06 (1.16)	−0.02 (−0.32)	0.01 (0.12)		−0.04 (−0.20)	0.03 (0.28)	0.06 (0.88)
D4	0.10 (0.80)	0.04 (0.65)	0.03 (0.50)		0.04 (0.52)	0.07 (1.76)	0.15** (3.47)		−0.04 (−0.20)	−0.06 (−0.93)	−0.04 (−0.59)
R-square	0.84	0.80	0.85		0.79	0.92	0.97		0.87	0.90	0.91
D. W.	1.96	2.24	2.01		2.20	2.18	2.44		2.57	2.56	2.60
AIC	−1.72	−2.91	−2.48		−0.62	−3.85	−4.57		−0.85	−3.24	−2.99
SIC	−1.13	−2.24	−1.84		0.15	−2.93	−3.43		0.12	−2.23	−2.18

[a] Number in parenthesis is asymptotic *t*-ratio. ** and * denote significance at the 5 percent and 10 percent levels, respectively. Since the model is estimated in double-log form, the coefficients of the exchange-rate and income variables are elasticities.

Table 3.5 *Wald* tests to determine whether bulk and high-value products can be aggregated

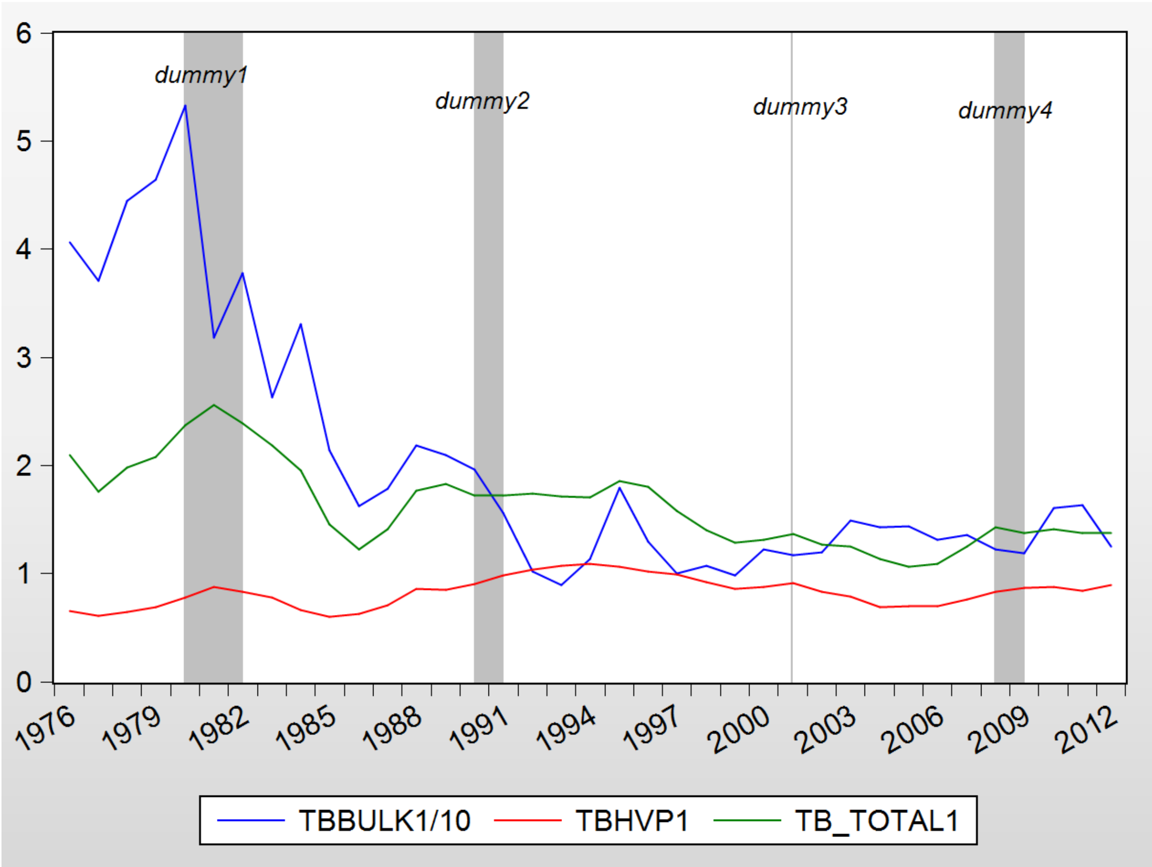
Model	Null Hypothesis	Computed *Chi-square*	Probability	Result
Export Value	Coefficients of Bulk and High Value Equations are Equal	127.4	0.00	Reject at 5% probability level
Import Value	Coefficients of Bulk and High Value Equations are Equal	101.9	0.00	Reject at 5% probability level
Trade Balance	Coefficients of Bulk and High Value Equations are Equal	165.7	0.00	Reject at 5% probability level

Table 3.6 *F*-tests to determine whether recession effects are jointly significant

Item	Computed *F*-Statistic[a]	Probability	Result
Export Value:			
Model A[b]	0.32	0.86	Fail to reject
Model B	1.39	0.29	Fail to reject
Model C	1.74	0.19	Fail to reject
Import Value:			
Model A	1.81	0.20	Fail to reject
Model B	1.28	0.35	Fail to reject
Model C	3.30	0.08	Reject at 10% probability level
Trade Balance:			
Model A	0.96	0.47	Fail to reject
Model B	1.40	0.30	Fail to reject
Model C	3.56	0.03	Reject at 5% probability level

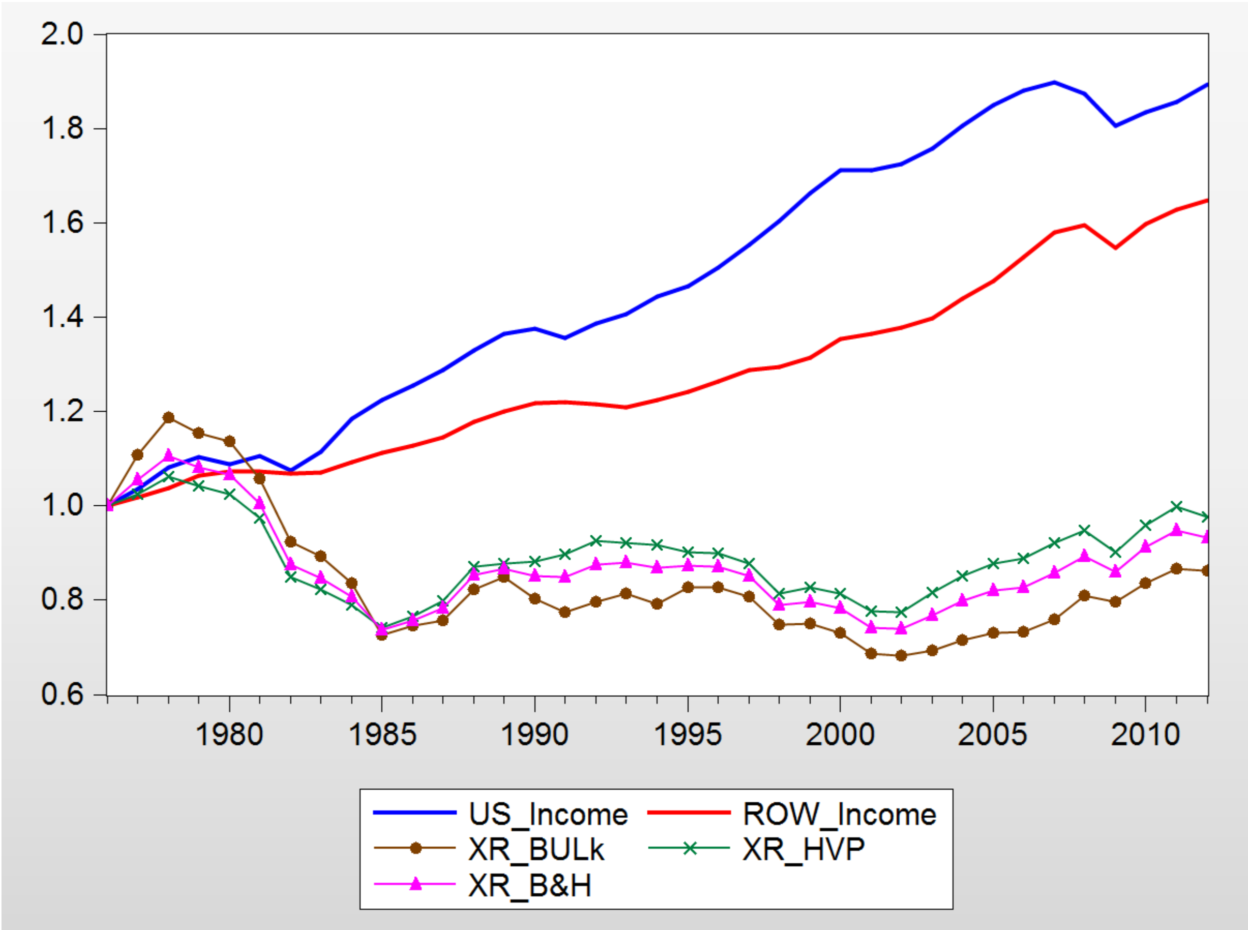
[a]The *F*-statistic is computed under null hypothesis that the coefficients of the dummy variables for the four recessionary periods are jointly zero.
[b]Models A, B, and C refer to bulk, high value, and combined bulk and high-value products, respectively.

Figure 3.1 Trade Balance Ratio for U.S. Bulk products and Real Trade Weighted Exchanges Rates for U.S. Bulk products, HVP and Combined agricultural goods. (2005=100)



Source: FAS, USDA

Figure 3.2 Domestic Income, Foreign Income and Real Trade Weighted Exchanges Rates for U.S. Bulk, High Value and the combined products respectively from 1976 to 2012 (1976=1).



Source: FAS, USDA, and ERS International Macroeconomic Data Set