THREE ESSAYS ON THE ECONOMICS OF VERTICAL AND SPATIAL RELATIONSHIPS IN SPECIALTY CROP SUPPLY CHAINS: CO2 EMISSION POLICIES, PRICE TRANSMISSION AND MARKET POWER

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Jun Lee
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THREE ESSAYS ON THE ECONOMICS OF VERTICAL AND SPATIAL RELATIONSHIPS IN SPECIALTY CROP SUPPLY CHAINS: CO2 EMISSION POLICIES, PRICE TRANSMISSION AND MARKET POWER

Jun Lee, Ph. D.
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This dissertation offers three essays addressing critical topics in supply chains for specialty crops: 1) impacts of initiatives to control CO2 emissions; 2) vertical price transmission behavioral changes after elimination of an export cartel; and 3) influence of market power on cost pass-through.

Chapter 1 investigates the impact of alternative CO2 emission reduction policies, namely a carbon tax and a cap-and-trade system, on the U.S. apple supply chain. The potential benefits accruing to farm-level CO2 sequestration options are also considered. A temporally- and spatially-disaggregated price equilibrium model is constructed to evaluate the relationship between CO2 emissions reduction and region- and sector-wide economic performance. The results indicate that all CO2 emission policies examined lead to modest decreases in CO2 emissions from the apple supply chain, with implied carbon prices ranging between $25 and $200 per metric ton of CO2. The results also suggest that a cap-and-trade system may be more cost-effective in reducing CO2 emissions than a carbon tax, regardless of the CO2 sequestration options utilized.

Chapter 2 examines the impact of terminating the coffee export quota system (EQS) on international-to-retail price transmissions in France, Germany and the United States. A threshold error correction model (TECM) is developed to measure price transmission behaviors, taking into account long-run threshold effects and short-run price transmission asymmetries (PTA). The results suggest that retail prices become more responsive to changes in international prices after the EQS elimination. The evidence suggests the presence of short-run PTAs, with significant differences across countries. We discuss these differences in terms of market structure.

Chapter 3 investigates links between exertion of market power and cost pass-through patterns in roasted coffee markets in the U.S. and Germany. A structural supply-demand model is developed to evaluate the degree of market power exerted by the roasting industry. Subsequently, a TECM is used to test the impacts of market power on cost pass-through behaviors. The results indicate that market power drives the existence of a "rockets and feathers" phenomenon in both countries. That is, market power causes retail prices to rise faster than they fall in response to changes in international prices.

BIOGRAPHICAL SKETCH

Jun Lee was born in Korea (Republic of) on October 13, 1975. His initial interest in economics emerged during the financial crisis that overwhelmed Korea in 1997. At the time, he was studying German Language and Literature at Korea University. However, in the rapid change of economic order and circumstance during the financial crisis, economics which was located in the center of disputation stimulated his academic need. He attended the Graduate School at Korea University where his research interests centered on inflation targeting policy and the estimation of core inflation rates. After receiving the M.S. degree in economic policy in 2003, he worked for two years as a researcher at the Korea Environment Institute (KEI) under the Office of the Prime Minister in Korea. The KEI is the leading national institute, dedicated to evaluating and consulting environmental policy and welfare in Korea. His research responsibilities at the KEI focused on sustainable development and environmental efficiency. Two years academic research experience enabled him to broaden his knowledge and understanding on environmental economic field as well as other topics in industrial organization field, and inspired him to study more advanced economic theory. He entered the Ph.D. program at the Dyson School of Applied Economics and Management, Cornell University in 2006. His Ph.D. research focused on several critical topics on environmental economics and industrial organization field. His research interests are alternative CO2 emissions reduction regulations and their impacts on food supply chain, vertical price transmission behaviors and influence of market power on cost pass-through in food manufacturing industry.

I dedicate this dissertation to my loving wife, Hye Jin and our precious son, Dong Ha. Certainly enough, am I indebted to them for their tolerance, perseverance, and every kind of support and devotion. There is no doubt in my mind that without their continuous support and sacrifice, I could not have completed my studies at Cornell University. I also send my deepest appreciation to my beloved parents for their emotional and spiritual support.

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The best and worst moments of my doctoral dissertation journey have been shared with many people. My enormous debt of gratitude can hardly be repaid to my good friends, Dr. Jae Sung Cho, a research fellow at Korea Rural Economic Institute and Dr. Jung Youn Mo, a research fellow at Korea Institute for Industrial Economics and Trade. Their support and care helped me overcome setbacks and stay focused on my graduate study. I greatly value their friendship and I deeply appreciate their belief in me. I am also thankful to my friend Julie Coulombe who helped me to improve my English and to keep a balanced perspective on life in Ithaca, New York. Most importantly, none of this would have been possible without the love and patience of my loving family. Their unwavering faith and confidence have sustained me throughout my Ph.D. life.

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CHAPTER 1

IMPACT OF CO2 EMISSION POLICIES ON FOOD SUPPLY CHAINS: AN APPLICATION TO THE U.S. APPLE SECTOR

Abstract

In this study, we investigate the impact of two alternative CO2 emission reduction policies, namely a carbon tax scheme and a cap-and-trade mechanism, on the economic performance of food supply chains. We also evaluate the potential impact of farm-level CO2 sequestration incentive programs on supply chain performance. For this purpose, we develop an optimization model for the U.S. fresh apple supply chain employing a spatially- and temporally-disaggregated price equilibrium approach. Our results suggest that, if carbon prices range between $25 and $200 per metric ton of CO2, all CO2 emission policies examined here lead to slight decreases in total CO2 emissions along the fresh apple supply chain. These results are primarily driven by decreases in apple production in supply regions. Specifically, we find that, given regional differences in price elasticities of supply, production costs and sequestration rates, California may experience the largest decreases in apple production with the introduction of CO2 emission reduction policies. The costs of CO2 emission policies affect primarily fresh apple consumers in the non-harvest season due to changes in apple storage usage. Farm-level carbon sequestration incentives seem to contribute modestly to CO2 emissions reduction, relative to policies without carbon sequestration incentives. We find that a cap-and-trade system may be more cost-efficient for CO2 emissions mitigation in the fresh apple supply chain than a carbon tax scheme, regardless of the presence of sequestration incentive programs.

1.1 Introduction

Tradeoffs between environmental quality improvement and economic performance have been widely addressed in the environmental economics literature. There is consensus that improving environmental quality entails substantial economic and social costs, but the social benefits obtained from pollution reductions often overwhelm these costs. A better understanding of these linkages is relevant for the design of cost-effective policies targeting environmental quality improvement.

Since the ratification of the Kyoto Protocol in 1997, many countries have introduced policy instruments that target reductions of carbon dioxide (CO2) emissions, a primary contributor to greenhouse gas (GHG) concentration. CO2 emissions are mostly generated by sectors where fossil fuels are heavily used, including electricity generation, transportation and industrial activities. These sectors have the largest share of CO2 emissions, accounting for 30 percent, 23 percent and 19 percent, respectively (US EPA 2011). In this regard, most policy instruments designed to reduce CO2 emissions have been implemented and discussed in the context of these three sectors.[1] We argue that efforts to regulate CO2 emissions can be extended to the food supply chain. The food supply chain consists of a variety of stakeholders including producers, processors, distributors and consumers, whose activities can influence the level of CO2 emissions at multiple segments.

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[1] For instance, the European Union's Emission Trading System (EU ETS), which is the largest international cap-and-trade program currently operating, has been applied to power stations, combustion plants, oil refineries and iron and steel works, as well as factories making cement, glass, lime, bricks, ceramics, pulp, paper and broad. In the United States, the Regional Greenhouse Gas Initiative (RGGI) in which seven U.S. states in the Northeast and Mid-Atlantic region participate, is a regional mandatory cap-and-trade program to control CO2 emissions from fossil fuel-fired electric generation.

The food supply chain accounts for a relatively small portion of CO2 emissions in the overall U.S. economy.[2] Nevertheless, there are many private- and public-led sector-specific initiatives to reduce CO2 emissions. Many industries are setting goals to reduce their carbon footprints to respond to increased consumer demand for sustainability attributed from food supply chain (Brenton, Edwards-Jones and Jensen 2010; Bockel et al. 2011). Therefore, from a sector-specific point of view, it is important for supply chain practitioners and policy makers to identify 1) the extent to which CO2 emission policies can contribute to emission reductions at each segment of the chain; 2) the extent to which the economic burden brought about by CO2 emission control policies is concentrated on particular segments of the chain; and 3) the extent to which CO2 emissions are offset by compensation for social costs. At the same time, particularly in food supply chains, it is also important for supply chain practitioners as well as policy makers to evaluate the potential economic impact of farm-level CO2 sequestration initiatives and their implications for long-term strategies to reduce CO2 emissions. In this study, we shed light on these issues by developing a spatially-disaggregated model for fresh apples in the United States and by evaluating the potential economic impacts of alternative CO2 emission reduction policies with and without carbon sequestration options.

The apple supply chain is an interesting and relevant one to examine the links between CO2 emissions reduction and economic performance. Table 1.1 presents selected characteristics of the U.S. apple sector. First, apples are one of the most popular fruits produced domestically, together with oranges and grapes. Specifically, apples are the most popular domestically produced fruit in fresh form and second to oranges in total fruit consumption. Second, the U.S.

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[2] The estimated food-related energy use which is closely related to CO2 emissions was estimated about 15.7 percent in the national energy budget in 2007 (Canning et al. 2010).

apple supply chain exhibits different consumption and production patterns spatially and temporally. According to the Continuing Survey of Food Intake by Individuals (USDA AMS 2011), among the four Census divisions, consumers in the West region have the highest consumption of fresh apples, whereas Northeast region consumers tend to eat more processed apples. In terms of total utilization, per capita apple consumption is highest in the Northeast and lowest in the South (Perez et al. 2001). Geographical differences in apple consumption may be related to the location of apple production regions. The six largest producer states (Washington, New York, Michigan, Pennsylvania, California and Virginia) account for approximately 90 percent of the country's total apple production (USDA NASS 2011). Particularly, Washington accounts for about 58 percent of total apple production in the country. The relatively high concentration of production in a few states and the dispersed pattern of apple consumption require the operation of multiple distribution networks in the apple supply chain. These characteristics provide meaningful insights on policy application for those who are interested in the relationship between supply chain performance and CO2 emissions reduction, in terms of 1) the extent of policy impact on each supply region; 2) the extent of policy impact on each chain segment; 3) the extent of policy impact on the distribution channel; 4) the extent of policy impact on consumers; and 5) the extent of policy impact on CO2 emission reductions and social welfare losses.

Table 1.1 Selected characteristics of the U.S. fresh apple sector

	Total	Fresh	Processed
Utilization[a, b]		*pounds*	
Oranges	65.2	7.5	57.7
Apples	49.9	16.4	33.5
Grapes	21.4	8.9	13.1
Demand[c]		*%*	
Northeast	29.1	24.7	30.9
Midwest	24.2	24.2	24.4
South	19.6	18.9	19.8
West	27.1	32.2	24.9
Supply[d]		*%*	
Northeast	15.6	31.7	21.1
Midwest	8.8	20.7	12.9
South	2.2	10.2	5.0
West	73.5	37.5	61.1

a. Source: Food availability per capita, U.S. Department of Agriculture, Economic Research Service.

b. Fresh equivalent, pound.

c. Source: Per capita apple consumption, CSFII 1994-1996 (1998), Perez et al. (2001).

d. Source: Noncitrus Fruits and Nuts 2010, U.S. Department of Agriculture, Economic Research Service.

Various approaches have been employed to measure environmental impacts associated with activities along the food supply chain. A common one is the Food Miles approach which measures total distances traveled by food from the farm gate to end consumers (Coley et al. 2009; Weber and Matthews 2009). The Food Miles concept is simple and easy to communicate to the public. That is, the longer food travels, the more energy is consumed through transport, so more GHGs are discharged to the atmosphere. However, the Food Miles approach does not consider economic and environmental impacts incurred by production, storage and processing activities, which account for the largest portion of CO2 emissions generated by a typical food supply chain (Nicholson, Gómez and Gao 2011). An alternative, and commonly used approach, is the Life Cycle Assessment (LCA) which covers a broader range of issues than the Food Miles approach. The LCA is a technique designed to evaluate environmental effects during the life span of products based on energy balance accounting. This approach was originally applied to manufactured products (Loss and Evans 2002). Several studies have extended LCA to food supply chains (Heller and Keoleian 2003; Andersson and Ohlsson 1999; Moller et al. 1996). However, LCA is a descriptive accounting approach that yields little or no insight into the potential impacts of CO2 emissions reduction on social costs and benefits as well as on the structure of the food supply chain. Consequently, a rigorous model showing the relationship between economic performance and CO2 emissions reduction in apple supply chain is required to identify appropriate policy instruments that may minimize the negative impacts of regulating CO2 emissions.

The objective of this study is to develop an optimization model of the U.S. apple supply chain to examine the impacts of alternative CO2 emission policies on social welfare and supply chain structure. We focus on two widely discussed policy instruments: a carbon tax scheme and a cap-and-trade program. For the purposes of this study, we employ a spatially- and temporally-disaggregated price equilibrium model, developed initially by Enke (1951) and Samuelson (1952), and later extended by Takayama and Judge (1964). Spatial price equilibrium models have been widely employed for the analysis of inter-regional competition of agricultural commodities (Guajardo and Elizondo 2003). These studies have refined the food supply chain model and examined how external factors such as supply shocks (Fuchs, Farris and Bohall 1974), transportation costs (Dunn and Garafola 1986), production costs (Chien and Epperson 1990), and policy changes (Yavuz et al. 1996; Guajardo and Elizondo 2003) influence supply chain structure and inter-regional competitiveness of commodities.

We make two key contributions to this extensive literature. First, we develop a spatially- and temporally-disaggregated model to examine the sector- and region-specific impacts of CO2 emission policies on the structure and performance of a food supply chain. This model provides specific insights that previous models were unable to find for both supply chain practitioners and policy makers interested in reducing CO2 emissions. Second, we extend the model by incorporating farm-level CO2 sequestration potential-a unique characteristic in agricultural product supply chain- and subsequently examine their potential impact on the food supply chain structure and their contribution to CO2 emission mitigation.

The remainder of this study is organized as follows. Section 2 describes the U.S. apple supply chain and discusses the scope of the proposed model. In section 3, we develop an optimization model of the U.S. fresh apple supply chain employing a spatial price equilibrium framework. In this section, we econometrically estimate price elasticities of supply and demand, calculate costs associated to supply chain activities, and calculate the CO2 emissions occurring at each segment of supply chain. Here, we also calibrate the model using 2006 data, which is used as a benchmark for comparison in the subsequent analysis. In section 4, we discuss CO2 emission policies and their application to our apple supply chain model. We also discuss the incorporation of farm-level carbon sequestration incentives in this section. Subsequently, we present four scenarios designed to examine the economic impacts and the effectiveness of alternative CO2 emission reduction policies on the fresh apple supply chain. In section 5, we discuss our simulation results. Our conclusions are elaborated upon in section 6.

1.2 The U.S. apple supply chain

Figure 1.1 illustrates the U.S. apple supply chain. Generally, apple orchards start the annual production cycle in early spring and apples are harvested in the fall. After harvest, apples enter either the fresh or the processed supply chain. Growers typically transport apples from the orchard to packing-shipping facilities, where apples are sorted for fresh or processed utilization and, subsequently, stored or shipped to processing facilities. Alternatively, some growers specialized in processed apples ship directly from the orchard to processing plants. Apples moved directly into the fresh supply chain are hand-picked and transported from the orchard to packing-shipping facilities. Apples are placed into one of two possible storage types. Apples for sales during harvest season (September to December) are put into regular (cold air) storage, whereas controlled atmosphere (CA) storage is used for fruit distributed during the non-harvest season (January to August). In both periods, fresh market apples are transported in trailer-trucks from packing sheds to retail distribution centers, which are generally located near consumption locations. Fresh apples are also traded in international markets. The United States imported fresh apples by 343.4 million pounds in 2009, which accounts for about 7 percent of annual demand.

Particularly, fresh apples imports peak primarily between April and June. U.S. fresh apples are also exported to other countries. In 2009, the United States exported 1,777.7 million pounds of fresh apples, which accounts for about 30% of total fresh apple production in the country.

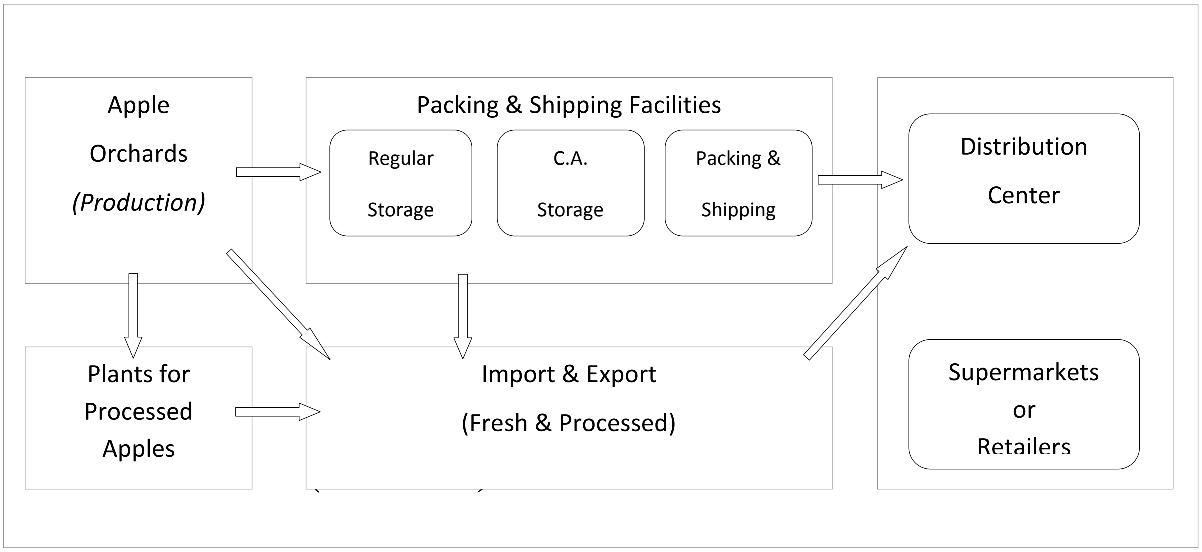

Figure 1.1 Diagram of the fresh apple supply chain

Our model focuses solely on the supply chain for fresh apples. The model considers six apple producer states (Washington, New York, Michigan, California, Pennsylvania and Virginia) and forty-nine consumption locations, each corresponding to the country's continental states. We take into account five fresh apple varieties (Red Delicious, Golden Delicious, Granny Smith, Gala and All others) to accommodate regional specialization on specific varieties (e.g. Granny Smith in Washington) and regional differences in consumer preference. The model is inter-temporal and, considers two time periods: (1) the harvest season (September to December), during which apples are primarily put into short-term storage prior to distribution; and (2) the non-harvest season (January to August), when apples are put into long-term CA storage. Our model takes into account exports and imports of fresh apples. The primary transportation method employed in our model is heavy-duty diesel trucks which account for approximately 95 percent of total apple transportation (USDA AMS 2011).

1.3 The model

1.3.1 Spatial price equilibrium model

We employ a spatially- and temporally-disaggregated price equilibrium model following Takayama and Judge (1964) to estimate optimal product flows from supply regions to consumption locations. The model solves a quadratic programming problem to maximize social surplus measured as the sum of consumers and producers surplus less total costs resulting from all activities within the supply chain (see Appendix 1.1 for an analytical presentation of the model). The objective function of the model can be expressed as:

$$
\begin{aligned}
Max \quad & \sum_{t}\sum_{k}\sum_{j}\delta^{t-1}\int_{0}^{qd^{t}_{k,j}} pd^{t}_{k,j}(qd^{t}_{k,j})dqd^{t}_{k,j} - \sum_{k}\sum_{i}\int_{0}^{qs_{k,i}} ps_{k,i}(qs_{k,i})dqs_{k,i} \\
& + \sum_{t}\delta^{t-1}\int_{0}^{qim^{t}} ppim^{t}(qim^{t})dqim^{t} - \sum_{t}\delta^{t-1}\int_{0}^{qex^{t}} ppex^{t}(qex^{t})dqex^{t} \\
& - \sum_{k}\sum_{i} pc_{k,i}\cdot qs_{k,i} - \sum_{t}\sum_{s}\sum_{i}\delta^{t-1}\left[sc^{t}_{s,i}\cdot\left\{\sum_{k} sa^{t}_{k,s,i}\right\}\right] \\
& - \sum_{t}\sum_{i}\sum_{j}\delta^{t-1}\left[tc^{t}_{i,j}\cdot\left\{\sum_{k}\sum_{s} tfa^{t}_{k,s,i,j}\right\}\right] - \sum_{t}\sum_{m}\sum_{j}\delta^{t-1}\left[tc^{t}_{m,j}\cdot\left\{\sum_{k} tfam^{t}_{k,m,j}\right\}\right]
\end{aligned}
\quad , \tag{1.1}
$$

where the objective function represents surpluses from both domestic and international sources of fresh apples minus total costs from all practices including production, storage and transportation. The model yields optimal quantities for production, consumption, exports and imports; producer prices for each supply region; retail prices for each consumption location; inter-regional commodity flows; and social welfare levels. The objective function in equation (1.1) is constrained to ensure balance between supply and demand, capacity restrictions (e.g. land capacity for production and the capacity for storage), and technical constraints in production and storage (e.g. yield rate in production and loss rate in storage). These constraints are stated as follows:

$$qs_{k,i} \le \mu_i \cdot PCAP_{k,i} \quad \forall k,i \text{, where } PCAP_{k,i} = aveyield_{k,i} \times \theta_{k,i} \cdot acre_i \quad \forall k \tag{1.2}$$

$$\sum_t \sum_s \frac{1}{\lambda^t_{s,i}} sa^t_{k,s,i} \le qs_{k,i} \quad \forall k,i \text{, where } sa^{t=spring}_{k,s=rs,i} = 0 \quad \forall k,i \tag{1.3}$$

$$\sum_t \sum_k sa^t_{k,s,i} \le SCAP_{s,i} \quad \forall s,i \tag{1.4}$$

$$\sum_j tfa^t_{k,s,i,j} + qex^t_{k,s,i} \le \frac{1}{\lambda^t_{s,i}} sa^t_{k,i,s} \quad \forall t,k,s,i \tag{1.5}$$

$$qd^t_{k,j} \le \sum_s \sum_i tfa^t_{k,s,i,j} + \sum_j tfam^t_{k,m,j} \quad \forall t,k,j \tag{1.6}$$

$$\sum_j tfam^t_{k,m,j} \le qim^t_{k,m} \quad \forall t,k,j \tag{1.7}$$

Equation (1.2) ensures that the total quantity of apple variety *k* produced in supply area *i*, $qs_{k,i}$, does not exceed the maximum production capacities, $\mu_i \cdot PCAP_{k,i}$, of each region, where μ_i is the percent rate of fresh utilization in supply region *i*. The maximum production capacities are determined by average yield, $aveyield_{k,i}$, production percentage of variety *k*, $\theta_{k,i}$ and total acreages in production region *i*, $acre_i$. Equations (1.3) and (1.4) specify the balance of apples moved from orchards to storage facilities and the maximum storage capacity in each production region *i*, respectively. In equation (1.3), the total amount of stored apples in each region *i*, $sa^t_{k,s,i}$, after considering storage loss rates, $\lambda^t_{s,i}$, must be less than or equal to total fresh apple supply in region *i*, $qs_{k,i}$. Following standard management practices, apples for sale during the harvest season are put into regular storage, whereas those for sale during the non-harvest season are put into CA storage. In the model, apples for sale in the harvest season can be distributed from either regular or CA storage. In contrast, apples for sales during the non-harvest season can be distributed only from CA storage. Therefore, the amounts stored in regular storage during the non-harvest season are set to equal to zero in our model. Equation (1.4) ensures that the total amount of stored apples at storage facility *s* in each supply region *i*, $sa^t_{k,s,i}$, must be less than or equal to the maximum storage capacity, $SCAP_{s,i}$. Equation (1.5) ensures that the total amount of apples in storage facility *s*, taking into account storage losses must be greater than or equal to the sum of shipments to both consumption locations, $tfa^t_{k,s,i,j}$, and export point of shipment, $qex^t_{k,s,i}$. Equation (1.6) ensures that fresh apple demand for variety *k* in location *j* at time *t*, $qd^t_{k,j}$, is less than or equal to the total shipments from all supply regions *i*, $tfa^t_{k,s,i,j}$, plus imports from all import points *m*, $tfam^t_{k,m,j}$. Equation (1.7) specifies that total shipment from import point *m* to consumption location *j*, $tfam^t_{k,m,j}$, must not exceed total imports at point *m* at *t*, $qim^t_{k,m}$. Finally, all decision variables in the model are required to be non-negative. Table 1.2 presents the indices and variables used in the model.

Table 1.2 Index, variables and definitions

Index (variable)	Definition
t	time period
k	apple variety
i	supply region
j	consumption location
s	storage facility
m	import and export port
$pd_{k,j}^{t}$	retail price of fresh apple variety *k* at time *t* in consumption location *j* ($/pound)
$qd_{k,j}^{t}$	fresh apple quantities consumed of variety *k* at time *t* in consumption location *j* (million pounds)
$ps_{k,i}$	price received by farmer for apple variety *k* in supply region *i* ($/pound)
$qs_{k,i}$	fresh apple quantities produced of apple variety *k* in supply region *i* (million pounds)
$ppim^{t}$	price differential between import price and retail price at time *t* ($/pound)
qim^{t}	quantities of imported apple at time *t* (million pounds)
$ppex^{t}$	price differential between export price and producers price at time *t* ($/pound)
qex^{t}	quantities of exported apple at time *t* (million pounds)
$pc_{k,i}$	production cost for apple variety *k* in supply region *i* ($/pound)
$sa_{k,s,i}^{t}$	fresh apples of variety *k* put into storage facility s at time *t* in region *i* (million pounds)
$sc_{k,s,i}^{t}$	storage cost to operate storage facility *s* at time *t* in region *i* ($/pound)
$tc_{i,j}^{t}$	transportation cost at time *t* from supply region *i* to consumption site *j* ($/pound)
$tc_{m,j}^{t}$	transportation cost at time *t* from supply region *i* to port *m* ($/pound)
$tfa_{k,s,i,j}^{t}$	shipment of apple variety *k* at time *t* from storage facility *s* in supply region *i* to consumption location *j* (million pounds)
$tfam_{k,m,j}^{t}$	shipment of apple variety *k* at time *t* from port *m* to consumption site *j* (million pounds)
$PCAP_{k,i}$	production capacity for variety *k* in supply region *i* (million pounds)
$aveyield_{k,i}$	average yield per acre in supply region *i* (pounds)
$\theta_{k,i}$	production percentage of variety *k* in supply region *i* (%)
$acre_{i}$	bearing acreage for apple in supply region *i* (acres)
$\lambda_{s,i}^{t}$	apple loss rate at storage *s* in supply region *i* at time *t* (%)
$SCAP_{s,i}$	capacity of storage facility *s* in supply region *i* (million pounds)
μ_{i}	percentage rate of fresh utilization in supply region *i* (%)

1.3.2 Estimation of price elasticities

The spatial equilibrium model employs price-dependent linear supply and demand functions to formulate the objective function. A linear inverse supply (demand) function can be written as:

$$p_{k,i(j)} = \alpha_{k,i(j)} + \beta_{k,i(j)} \cdot q_{k,i(j)}, \tag{1.8}$$

where *k* is the apple variety, *i* (*j*) is the supply (demand) region. The coefficient $\beta_{k,i(j)}$ is obtained from the relationship $\beta_{k,i(j)} = \frac{1}{\varepsilon_{k,i(j)}} \cdot \frac{p_{k,i(j)}}{q_{k,i(j)}}$, where $\varepsilon_{k,i(j)}$ refer to price elasticities of supply (demand) in region *i* (*j*). Therefore, estimating price elasticities of supply (demand) is required to establish the linear relationship between prices and quantities in supply and demand.

We employ the Linear Approximate Almost Ideal Demand System (LA-AIDS), developed by Deaton and Muellbauer (1980) to estimate price elasticities of demand for each variety, time period and demand location. The LA-AIDS is useful for capturing consumer demand patterns in household-level micro data by controlling for different socio-economic variables (e.g. race, education level and gender) and income levels (Blundell, Pashardes and Weber 1993). We use Nielsen's Homescan panel database for the period 2005 and 2006 to estimate demand elasticities of fresh apples.[3]

While the use of household-level data is preferable in the complete demand system, one has to solve the zero expenditure problem presented in the data. That is, households do not always purchase all apple varieties in each time period specified in the model. If this problem is

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[3] The households reporting fresh apple purchase in 2005 and 2006 numbered 5,642 and 4,736, respectively. The sample size used in the estimation is 20,297.

not addressed, parameter estimates can be biased (Park et al. 1996). We follow the censored method of Heckman (1978) to deal with the zero consumption problem. First, we estimate a probit model to determine the probability that a given household would purchase a specific variety *k* as follows:

$$q_{k,h} = f(p_{1,h},\cdots,p_{k,h},X_h,\eta_{1,h},\cdots,\eta_{N,h}), \tag{1.9}$$

where $q_{k,h}$ is a binary variable equal to 1 if variety *k* is purchased by household *h*, zero otherwise; $p_{k,h}$ is the price of variety k ; $\eta_{n,h}$ are the socio-economic variables describing household characteristics ($n=1,\cdots,N$); and $X_{t,h}$ is the expenditure level. Second, using the fitted value, $\hat{q}_{k,h}$, from the probit estimation in equation (1.9), we compute the inverse Mills ratio, $MR_{k,h} = \phi(p_h,\eta_h,X_h)/\Phi(p_h,\eta_h,X_h)$ for households consuming apple variety *k*. $\phi(\cdot)$ and $\Phi(\cdot)$ refer to the density and cumulative probability functions, respectively (Heien and Wessells 1990). Similarly, $MR_{k,h} = \phi(p_h,\eta_k,X)/\{1-\Phi(p_k,\eta_k,X)\}$ is computed for households that do not purchase the variety *k*. Finally, we estimate the missing prices stemming from zero purchase of variety *k* at time *t*, employing seasonal and regional dummy variables for the complete data set. Consequently, with the complete data set, the LA-AIDS model incorporating the inverse Mills ratio, MR_k is given by:

$$w_{k,h} = \sigma_k + \sum_n \varsigma_{k,n}\eta_{n,h} + \sum_l \gamma_{k,l} \ln p_{l,h} + \pi_k \ln(X_h / P_h) + \delta_k MR_{k,h} + \rho T , \tag{1.10}$$

where $w_{k,h}$ is the budget share of household *h* for variety *k*; $X_h = \sum_{k=1}^{5} p_k q_{k,h}$ is the total expenditure of household *h* for five apple varieties; $MR_{k,h}$ is the inverse Mills ratio; $\ln P_h = \sum_{k=1}^{5} w_k \ln p_{k,h}$ is the Stone Price Index for the LA-AIDS specification; and T is year dummy variable. To be consistent with economic theory, the parameter restrictions for estimation of equation (1.10) include the following: for homogeneity, $\sum_{k=1}^{5} \sigma_k = 1$, $\sum_{k=1}^{5} \pi_k = 0$, $\sum_{k=1}^{5} \gamma_{k,l} = 0$, $\sum_{l=1}^{5} \gamma_{k,l} = 0$; and for symmetry, $\gamma_{k,l} = \gamma_{l,k}$. Finally, the own price elasticity of demand is obtained by $\hat{\varepsilon}_{kk} = \frac{\gamma_{kk} - \pi_k w_k}{w_k} - 1$.

Table 1.3 reports the estimated seasonal price elasticities of demand for four multistate Census regions, varieties and seasons, respectively (See Appendix 1.3 for the detailed estimation results of the LA-AIDS model in equation (1.10) for the four Census divisions). The results show differences in the price elasticities of demand across regions, varieties and seasons. Red Delicious apples tend to have the same elasticity, regardless of the region, but the elasticity is higher in the non-harvest season. The other varieties (Golden Delicious, Granny Smith and Gala) exhibit elasticities that appear to be more sensitive to changes in prices compared to Red Delicious, the most consumed variety.

Table 1.3 Price elasticities of demand by season, region and variety

Time period	Region			
Harvest season (September-December)	Northeast	Midwest	South	West
Golden Delicious	-1.537	-1.168	-0.970	-0.608
Granny Smith	-3.349	-1.490	-2.000	-2.078
Red Delicious	-0.984	-1.016	-0.998	-0.933
Gala	-1.517	-0.691	-0.787	-1.120
Others	-1.054	-1.079	-1.080	-1.087
Non-harvest season (January-August)	Northeast	Midwest	South	West
Golden Delicious	-2.004	-2.712	-1.710	-3.222
Granny Smith	-2.562	-4.681	-1.955	-2.693
Red Delicious	-1.002	-1.105	-0.989	-0.901
Gala	-0.707	-1.269	-0.717	-0.959
Others	-1.057	-1.077	-1.098	-1.063

We also need to estimate price elasticities of supply in the production regions. For this purpose, we employ Nerlove's (1956) model and yearly data showing fresh apple production and farm gate prices for the period 1973 to 2008. Since these are national data, price elasticities of supply are assumed to be identical across all varieties in each supply region. Following Nerlove (1956), the general supply function of partial adjustment is specified as follows:

$$\ln Q_t^{i*} = \alpha_0 + \alpha_1 \ln p_{t-1}^i + \alpha_2 z^i + \alpha_3 T + u_t^i , \tag{1.11}$$

where Q_t^{i*} implies desired (equilibrium) output in production region *i* at time *t*, $\ln p_{t-1}^i$ is output price in region *i* at time *t-1*, T is a time trend and z^i represents weather conditions (e.g. precipitation). The supply dynamics for the actual adjustment to the equilibrium relationship is given by:

$$\ln Q_t^i - \ln Q_{t-1}^i = \varphi^i (\ln Q_t^{i*} - \ln Q_{t-1}^i) , \tag{1.12}$$

where Q_t^i is the actual output in region *i* and φ^i is the partial adjustment coefficient. By substituting equation (1.12) into equation (1.11), rearranging terms and employing regional dummy variables, the Nerlovian model for estimating price elasticities of supply for each supply region is given as:

$$\ln Q_t^i = \beta_0 + \beta_1 \ln p_{t-1}^i + \sum_j \beta_1^j \cdot d^j \cdot \ln p_{t-1}^i + \beta_2 \ln Q_{t-1}^i + \beta_3 z^i + \beta_4 T + e_t^i , \tag{1.13}$$

where $d^j = 1$ if $i = j$ and zero otherwise. $\beta_1 = \varphi \alpha_1$ is the short-run price elasticity of apple supply in each region. Table 1.4 presents the estimation results of equation (1.13). Based on the estimates, the price elasticities of supply are obtained for each supply region. The estimated elasticities for each supply region suggest that supply in California tends to be the most sensitive with respect to farm gate prices among all six production regions. In contrast, supply in Washington seems to be the most inelastic to farm gate prices.

Table 1.4 Estimate results of Nerlovian model and estimated price elasticities of supply

	Coefficient	Standard Errors
Constant	2.818**[a]	(1.199)
lnp_{t-1}	0.359**	(0.105)
$d_{CA} \cdot lnp_{t-1}$	0.206**	(0.069)
$d_{MI} \cdot lnp_{t-1}$	0.012	(0.040)
$d_{PA} \cdot lnp_{t-1}$	0.143**	(0.059)
$d_{VA} \cdot lnp_{t-1}$	0.186**	(0.068)
$d_{WA} \cdot lnp_{t-1}$	-0.242**	(0.068)
lnq_{t-1}	0.746**	(0.045)
Precipitation	-0.020	(0.029)
Temperature	-0.006	(0.017)
Time	-0.009**	(0.003)
R^2	0.95	
# of Observations	210	
Price Elasticities of Supply		
California	0.565	
Michigan	0.359	
New York	0.359	
Pennsylvania	0.501	
Virginia	0.545	
Washington	0.117	

a. ** indicates 5% significance level.

Finally, we estimate price elasticities of apple imports and exports for the harvest and non-harvest seasons. The quantities of exports and imports are assumed to be dependent on the differential between domestic and international import (export) prices and domestic production quantities.[4]

1.3.3 Costs in production, storage and transportation

To build the optimization model in equation (1.1), a specification of cost structure in the fresh apple supply chain is required. To this end, we compiled per unit costs for production, storage and transportation activities. Apple production costs consist of variable costs (e.g. labor, fertilizer, agricultural chemical) and ownership costs (e.g. depreciation, interest, insurance and tax). Labor costs are primary input during apple production, requiring when pruning, thinning, and apple picking. Moreover, fertilizer and agricultural chemical costs also account for a substantial part in apple production cost. Variable costs in production account for about 63 percent of total apple production costs. In contrast, ownership costs refer to overhead cost such as depreciation, interest, insurance and tax. In general, ownership costs in apple production account for about 37 percent of total apple production costs (Schotzko and Granatstein 2005). We compiled production costs from existing apple crop budget studies and generated the other states' costs by adjusting general farm unit labor costs in each state (BLS 2011).[5] Storage costs depend on the amount of time that apples are stored and the costs of energy. Storage costs were obtained from a direct survey of

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[4] Due to data limitations, we assume that fresh apples are exported only from Washington and New York. Furthermore, fresh apples are assumed to be imported through four major ports (e.g. LA, Miami, Newark and Seattle) and distributed to forty-nine consumption locations in the same proportion. Linear import functions for the non-harvest season and the harvest season are estimated to be $p_{nh}=1.151+0.003im_{nh}$, $p_h=1.151+0.022im_h$, respectively. Linear export functions for non-harvest and harvest season are estimated to be $p_{nh}=6.423-0.005ex_{nh}$, $p_h=6.423-0.008ex_h$, respectively.

[5] See the Appendix 1.3 for data table for production costs.

apple packer-shippers,[6] and other states' data are generated by adjusting the industry's electricity price data for each state (BLS 2011). Transportation costs were compiled from the "Agricultural Refrigerated Truck Quarterly"(USDA AMS 2011) and missing data were generated taking into account the relationship between costs and distances.[7]

1.3.4 CO2 emissions in production, storage and transportation

Bringing apples from the orchard to consumers requires various activities including farm-level production, storage and transportation, which generate CO2 emissions released into the atmosphere. Therefore, accounting for CO2 emissions along the fresh apple supply chain is critical to assess the economic impacts of CO2 emissions reduction policies.

First, farm-level activities in orchards contribute to CO2 emissions in two ways: direct and indirect energy use. CO2 emissions from direct energy use are primarily associated with water irrigation which is a critical input in apple production in some parts of the United States. Because of regional differences in rainfall patterns, irrigation requirements are different across the supply locations. Table 1.5 shows CO2 emissions resulting from irrigation activities in each supply region (NYSERDA 2008). These data show that the Western states, California and Washington, are arid and require relatively more irrigation due to lower precipitation. In contrast, the Eastern states need less energy for irrigation in apple production. However, given differences in yields across supply regions, CO2 emissions per million pounds of apples are lowest in Washington and highest in California.

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[6] We collected storage costs data from H.H. Dobbins, Inc in Lyndonville, NY. In western New York, the storage costs operating regular (cold air) storage are $23.50 per bin. In contrast, CA storage costs $33.50 per bin for apple storage. See the Appendix 1.3 for data table for storage costs.

[7] See the Appendix 1.3 for data table for transportation costs.

Table 1.5 CO2 emissions per unit apple production and storage

	CA	MI	NY	PA	VA	WA
	Metric Ton of CO2 per million pounds					
Production						
Direct use[a]						
Rainfall (*inch*)	0.5	3.0	3.0	3.5	3.5	1.0
Energy use (*kBTU/acre*)	20,347	3,393	3,391	3,393	3,393	23,739
CO2 emissions	312.98	208.65	163.29	185.97	185.97	154.22
Indirect use[b]						
CO2 emissions from fertilizer use	3.145	2.367	1.892	1.197	2.783	2.328
CO2 emissions from ag-chemical use	17.246	10.770	8.198	8.629	15.699	11.943
	Metric Ton of CO2 Ton per Acre					
CO2 sequestration rate[c] during apple production	0.105	0.145	0.129	0.129	0.129	0.105
Storage[d]	*Metric Ton of CO2 per million pounds*					
Regular Storage	80.45					
CA Storage	314.45					
Transportation[e]	*Kg Ton of CO2 per mile per truck*					
Heavy-diesel truck (class 8)	1.64					

a. Source: NYSERDA (2008).

b. Source: Saunders, Barber and Taylor (2006).

c. Source: Kroodsma and Field (2006).

d. Source: Canals et al. (2004).

e. Source: Mobile 6.2, US EPA (2011)

Another CO2 emission source in farm-level apple production comes indirectly due to the application of fertilizer and agricultural chemicals.[8] Data on fertilizer and agricultural chemicals usage in each supply location are from the U.S. Department of Agriculture (USDA NASS 2008). Subsequently, the data were converted into CO2 emissions following Saunders, Barber and Taylor (2006). Table 1.5 reports CO2 emissions resulting from the use of fertilizers and agricultural chemicals. Our accounting process suggests that, on a per unit basis, CO2 emissions are highest in California. In contrast, Pennsylvania shows the lowest emissions in terms of fertilizer and agricultural chemical use.

Although food supply chain activities release CO2 to the atmosphere, agricultural practices can capture a considerable amount of CO2 in soils and biomass (Lewandrowski et al. 2004). The extent to which carbon is sequestered varies between crop types. Particularly, orchards extract relatively more CO2 from the atmosphere than the average of all crops. For instance, non-silage annual crop capture an average of 14 g of CO2 per m^2 annually, whereas orchards sequester 26 g of CO2 per m^2 annually (Kroodsma and Field 2006). We take into account CO2 capturing, naturally occurring in orchard farms. Table 1.5 reports annual carbon sequestration rates during apple production by each region. Given differences in soil characteristics and climatic conditions across supply regions, carbon sequestration rates are relatively low in the Western states (California and Washington) and is relatively high in the Eastern states (New York, Pennsylvania and Virginia). Net CO2 emissions in apple production are obtained by deducting the sequestered CO2.

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[8] Components of fertilizer such as nitrogen, phosphate, potash and sulfur are converted to CO2 when they are applied to soil. Furthermore, the use of agricultural chemicals including herbicide, insecticide and fungicide also generates CO2 emissions.

Storage activities also discharge a substantial amount of CO2 into the atmosphere. As described in section 2, apples for fresh utilization are put into either regular or CA storage depending on the time of year that the apples are sold. Each storage facility is operated using electricity to maintain the proper temperature for apple freshness (Canals et al. 2007). Table 1.5 shows CO2 emissions generated by storage operation. Transport activities also emit CO2 during apple distribution. Fresh apples are mostly moved by heavy-duty trucks (class 8) which haul about 40,000 pounds with fuel economy of 6.1 mpg (King et al. 2010)[9]. We compiled CO2 emission data in fresh apple transportation using the MOBILE 6.2 model developed by the U.S. Environmental Protection Agency (US EPA 2011). We used the average CO2 emission rate of 1.64 kg/mile with fuel economy of 6.1 mpg.

1.3.5 Model calibration and baseline solutions

Before calibration, we constructed the linear supply and demand functions as in equation (1.8) using the estimated price elasticities of supply and demand, quantities for supply and demand, and producer prices and retail prices, respectively. We calibrated the model using the General Algebraic Modeling System (GAMS). The estimated solutions were compared with actual data in 2006 for model verification. Table 1.6 presents the baseline solutions of the model, the actual 2006 data and the ratio of our estimated to the actual value for all supply locations and for selected consumption locations.

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[9] There are three transportation options depending on distribution networks: 1) a semi-trailer hauling 40,000 pounds of apples; 2) a mid-size truck hauling 10,000 pounds of apples; and 3) a pick-up truck hauling 1,000 pounds of apples (King et al. 2010).

Table 1.6 Baseline solutions and comparison with actual data in 2006

	Estimated[a]	Actual	Estimated/Actual (%)
Total Supply (*million pounds*)	**5,523**	**5,656**	**98**
Quantity (*million pounds*)			
California	146	155	94
Michigan	286	295	97
New York	678	690	98
Pennsylvania	129	132	98
Virginia	32	34	94
Washington	4,251	4350	98
Producer Price ($/pound)			
California	0.370	0.412	90
Michigan	0.222	0.245	91
New York	0.289	0.302	96
Pennsylvania	0.247	0.257	96
Virginia	0.218	0.238	92
Washington	0.274	0.313	88
Export (*million pounds*)	1,495	1,508	99
Import (*million pounds*)	376	345	109
Total demand (*million pounds*)	**4,454**	**4,493**	**99**
Retail Price (Non-harvest) ($/pound)			
Atlanta	1.37	1.25	110
Los Angeles	1.31	1.16	113
New York	1.37	1.24	110
Philadelphia	1.37	1.26	109
Retail Price (Harvest) ($/pound)			
Atlanta	1.33	1.40	95
Los Angeles	1.28	1.27	101
New York	1.34	1.33	101
Philadelphia	1.34	1.34	100

a. Source: Authors' calibration

The estimated total production accounts for around 98 percent of actual apples produced in all six states. Due to the slightly low quantities estimated by the model, prices in each supply region are also modestly underestimated in comparison to actual producer prices in 2006. These differences may be because we do not take into account direct distribution systems (e.g. orchard to retailers and orchard to consumers). The estimated retail prices for the harvest season are practically the same as the actual retail prices in selected consumption locations (Table 1.6). However, the estimates for the non-harvest season are higher in comparison to actual data. These differences may be due to under-estimated production quantities, leading to low supplies of fresh apples in the non-harvest season in the model.

Table 1.7 presents CO2 emissions derived from the baseline model. The results show that the largest emissions along the fresh apple supply chain are attributable to farm-level apple production activities, accounting for about 46.6 percent of total supply chain CO2 emissions. Storage activities rank second in CO2 emissions, generating 43.6 percent of total CO2 emissions. Emissions from transportation activities are relatively low compared to production and storage activities, discharging 9.8 percent of total CO2 emissions in supply chain.

Table 1.7 CO2 emissions along the fresh apple supply chain

	CA	MI	NY	PA	VA	WA	Total	%
				Metric Ton of CO2				
Production[a]	48,613	63,360	117,631	25,228	6,633	716,300	977,764	46.6
Regular storage[a]	10,196	13,553	37,070	3,855	408	150,480	215,563	10.3
CA storage[a]		31,130	55,760	22,551	7,765	582,075	699,281	33.3
Transportation[a]							206,375	9.8
Total Emission							2,098,983	100

a. Source: Authors' calibration

1.4 Empirical framework for CO2 emission policies

1.4.1 CO2 emission policies and application to the model

Policy instruments to address CO2 emission reductions can be characterized into two categories. The first category is an incentive-based policy such as a carbon tax scheme or a cap-and-trade system. The second category is a regulatory-based policy which includes a non-tradable cap and a technology or a performance standard. In this study, we consider two alternative incentive-based policy instruments targeting CO2 emission reductions: a carbon tax scheme and a cap-and-trade program. A carbon tax scheme levies a tax on a per-unit amount of CO2 released into the atmosphere. In contrast, a cap-and-trade program sets upper limits on total CO2 emissions within a certain industry or a certain region, and gives the participants the right to trade CO2 emission permits. Given the non-point source nature of CO2 emissions from transportation activities, we apply these policy instruments only to CO2 emissions generated from apple production and storage activities.

To implement a carbon tax scheme in our model, let τ be a carbon tax rate levied on a metric ton of CO2 (mt C) emitted from apple production and storage activities. It is important to note that a carbon tax is not an *ad valorem* tax, but a specific (per unit) tax (Goulder 1992). Therefore, a carbon tax raises the production and storage costs proportionally to tax rates and to CO2 emission rates. In our model, in equation (1.1), the new production cost function that orchard owners face becomes $p\hat{c}_{k,i} = pc_{k,i} + \tau \cdot \mu_{p,i}$, where $\mu_{p,i}$ represents the CO2 emission rate per unit of apples production in supply region *i*. Similarly, the new storage cost function becomes $s\hat{c}^{t}_{s,i} = sc^{t}_{s,i} + \tau \cdot \mu^{t}_{s,i}$, where $\mu^{t}_{s,i}$ refers to the CO2 emission rate per unit of apples put into storage facility *s* at time *t* in region *i* (see Appendix 1.2 for a carbon tax policy application to the baseline model).

We now consider a market-based emissions cap-and-trade program. Let E_i represent CO2 emissions from supply region *i*, where total emissions are $\overline{E} = \sum_i E_i$. Let A_i be the lump-sum allocation of emission permits to the supply region *i* where total emission permits are exogenously determined by the emission reduction plan. That is, $\sum_i A_i = (1-\psi)\overline{E}$, where ψ is the predetermined emission reduction plan. Let ω refer to the permit price. To implement a cap-and-trade program in this analysis, the objective function in equation (1.1) is modified by adding the term, $\sum_i \omega \cdot [A_i - \mu_{p,i} \sum_k qs_{k,i} - \mu^{t}_{s,i} \sum_t \sum_k sa^{t}_{k,s,i}]$, which measures the benefits from selling or buying emission permits to or from other regions and segments of fresh apple supply chain. In a cap-and-trade system, emission permits can be distributed to each entity by grandfathering or auctioning. Under a grandfathering, permit allocation in each *i*, A_i is freely distributed based on the predetermined allocation rule (e.g. output-based or performance-based rule). In contrast, under auctioning, $A_i = 0$. In addition, total emissions are controlled by a new set of constraints, $\sum_i [\mu_{p,i} \sum_i qs_{k,i} + \mu^t_{s,i} \sum_t \sum_k sa^t_{k,s,i}] \leq (1-\psi)\overline{E}$, which ensures that total CO2 emissions from production and storage activities must be less than or equal to an exogenously determined emission reduction plan (see Appendix 1.2 for a cap-and-trade program policy application to the baseline model).

Furthermore, our model takes into account the potential value of carbon sequestration in farm-level production. As one of the important methods of mitigating CO2 emissions, carbon sequestration provides an additional incentive for agricultural producers who might participate in this program by switching production practices or land use. We consider two carbon sequestration options: (i) changes in production practices from conventional orchard operation to no orchard operation; (ii) changes in land use from orchard to forest. Carbon sequestration rates differ across regions because they depend on various local factors such as climatic conditions, soil characteristics, historical land use patterns and current management practices (Lewandrowski et al. 2004). Table 1.8 shows data suggesting that sequestration rates for changes in land use from orchard to forest are higher than those for changes in production practice from conventional orchard operation to no orchard operation.[10] However, converting orchard into forest is accompanied by a site preparation and tree establishment costs.[11] Table 1.8 reports the estimated costs for forest establishment. If the policy makers introduce carbon sequestration

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[10] To obtain the estimated average annual carbon sequestration rate, first, the highest value for CO2 accumulated in ecosystem for each of the first 15 years is calculated. Then, these values are divided by 15.

[11] General forest-establishment costs are not available for most regions. Lewandrowscki et al. (2004) generated estimates of forest establishment costs for each region following DuBois et al. (1999) and USDA FS (1999).

incentive programs, orchard owners may benefit from payments for offset credits created by the following carbon sequestration practices: i) no orchard operation and ii) converting their orchard into forest. Carbon sequestration is formulated in our model by adding a term which describes benefits from changes in land use or changes in production practices (see Appendix 1.2 for carbon sequestration incentive programs application to the baseline model).

Table 1.8 Selected carbon sequestration rates and forest establishment costs

State	From conventional operation to no operation[a] (*Metric Ton/Acre*)	From orchard to forest[a] (*Metric Ton/Acre*)	Forest establishment cost[b] (*$/Acre*)
California	0.120	0.786	108.93
Michigan	0.165	1.331	75.85
New York	0.147	1.210	75.85
Pennsylvania	0.147	1.210	75.85
Virginia	0.147	1.573	123.51
Washington	0.120	0.786	108.93

a. Source: Lewandrowski et al. (2004).

b. Source: Birdsey (1996); Lewandrowski et al. (2004).

1.4.2 Policy scenarios

Based on the values of the economic variables and CO2 emissions derived from the baseline solution, we explore the impacts of alternative CO2 emission policies on the U.S fresh apple supply chain. Specifically, we focus on 1) the extent of contribution to CO2 emission reduction; 2) the extent of changes in fresh apple supply chain structure; 3) the potential impact of carbon sequestration incentives; and 4) the cost effectiveness of alternative policies.

We examine four alternative policy scenarios in this study. In our simulation scenarios, carbon prices per mt C are exogenously determined. To trace out the marginal social costs for emission reductions, we run each scenario with eight alternative carbon price levels where the value of a mt C is $25, $50, $75, $100, $125, $150, $175 and $200, respectively.[12] Given that our model is based on a partial equilibrium framework, we do not consider general equilibrium issues that may take place in a multi-sector model and a second-best environment where distortionary taxes such as income, payroll and sales taxes exist. In this regard, we ignore the tax interaction effect and the revenue recycling effect.[13]

Scenario I examines the impacts of a carbon tax on the fresh apple supply chain. For this scenario, we assume that a social planner has perfect information on CO2 emission rates in this industry. We impose a tax per mt C generated by production and storage activities. In scenario I, various tax rates from $25 to $200 per mt C are imposed and examined, respectively. Scenario II also employs a carbon tax scheme, but additionally offers two possible carbon sequestration incentives to orchard owners. The first option (i) refers to an incentive program based on the sequestered CO2 resulting from changes in production practices from conventional orchard

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[12] Theoretically, optimal carbon prices are determined by the social cost of carbon (SCC) which is a marginal damage cost by CO2 emission. Various ranges of the SCC have been estimated through extensive literatures. According to Toll (2008), the SCC varied from $2 to $1677 per metric ton of CO2 in literatures. In this study, we employ the carbon prices range between $25 and $200 per mt C, frequently addressed in several studies (Bovenberg and Goulder 1996; Lewandrowski et al. 2004; Fischer and Fox 2007; Metcalf 2009).

[13] For example, in a general equilibrium framework, the imposition of a new tax (e.g. a carbon tax) on the economy raises output prices, leading to decreases in real wages, resulting in an excess burden to consumers. That is, the gross costs resulting from the imposition of a carbon tax are determined by pre-existing distortionary taxes which are already imposed in other sectors. On the other hand, gross costs of a carbon tax would be reduced by recycling the revenues from imposing a tax through lowering the distortionary taxes. (Goulder 1994; Fischer and Fox 2007).

operation to no orchard operation. The second option (ii) represents an incentive program targeting on the sequestered CO2 resulting from changes in land use from orchard to forest. Carbon payment levels for CO2 sequestration and carbon tax rates are equally set from $25 to $200 per mt C, respectively. In Scenario I and II, we do not consider the revenue recycling. That is, the revenues from newly imposing tax scheme are assumed to be ignored.

Scenario III simulates the potential effects of a cap-and-trade program on the fresh apple supply chain. In this scenario, emission permits are assumed to be allocated through auctioning, but the revenues from selling these permits are assumed to be ignored as in Scenario I and II. In addition, we allow the practitioners in the fresh apple supply chain to trade the permits with both the entities in supply chain and the ones outside supply chain. Each supply region would purchase CO2 emission permits to the extent which they would want to produce apples. Therefore, if CO2 emissions in a certain supply region are greater than the established baseline CO2 emissions, this region is considered to act as a net buyer of emission permits. In contrast, a region acts as net seller when CO2 emissions are less than the baseline emissions. For comparison with a carbon tax scheme, the emission cap is set by the total emission from production and storage activities in the baseline. Each region (entity) is assumed to be a price taker in the emission permit market. Lastly, Scenario IV incorporates carbon sequestration incentives with a cap-and-trade system. Like Scenario II, two sequestration incentive options (e.g. changes in production practice and changes in land use) are examined. Moreover, carbon payment levels for the sequestered CO2 and permit prices are also set equally in this scenario.

1.5 Simulation results

Empirical results are obtained from comparing the baseline model solution to the four alternative policy scenario simulation. Tables 1.9-1.12 and Figures 1.2-1.4 summarize selected results from each scenario, focusing on changes in economic variables and on CO2 emissions in the fresh apple supply chain. Simulation results should be interpreted as differences of each policy scenario relative to the baseline solution.

1.5.1 Carbon tax

Table 1.9 reports the percent change in outputs and prices for tax rates of $25, $100 and $200 per mt C under Scenario I. The results show that the imposition of a carbon tax of $25, $100 and $200 per mt C leads to decreases in total supply by 0.05, 0.15 and 0.28 percent, respectively. This suggests that a carbon tax in the range of $25 to $200 per mt C has modest effects on total fresh apple production because CO2 emission rates on a per unit basis are very small. For instance, a carbon tax of $25 per mt C raises the per-pound production costs by $0.0043-$0.0083 across supply regions. However, because of differences in CO2 emission rates and per unit apple production costs across supply regions, our simulation results show regional differences regarding the impact of a carbon tax. For example, California seems to be more responsive to a carbon tax policy than other supply regions. This is explained by the fact that California has the highest CO2 emission rates in apple production and the highest production costs among all regions.

Table 1.9 Percent changes (%) in output and price under a carbon tax scheme

	Scenario I			Scenario II					
	Carbon Tax			Carbon Tax with Sequestration Option I			Carbon Tax with Sequestration Option II		
Tax ratess	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**
Total Supply	**-0.05**	**-0.15**	**-0.28**	**-0.05**	**-0.15**	**-0.29**	**-0.03**	**-0.15**	**-0.30**
Quantity									
CA	-0.78	-2.23	-4.20	-0.79	-2.27	-4.28	-0.57	-2.12	-4.37
MI	-0.03	-0.08	-0.15	-0.03	-0.08	-0.16	-0.02	-0.09	-0.18
NY	-0.04	-0.07	-0.12	-0.04	-0.07	-0.12	-0.04	-0.08	-0.13
PA	-0.04	-0.10	-0.17	-0.04	-0.10	-0.18	-0.03	-0.10	-0.20
VA	-0.02	-0.09	-0.19	-0.02	-0.09	-0.20	0.00	-0.10	-0.24
WA	-0.02	-0.10	-0.19	-0.02	-0.10	-0.19	-0.01	-0.09	-0.20
Producer Price									
CA	-1.62	-4.32	-7.84	-1.62	-4.32	-8.11	-1.08	-4.05	-8.11
MI	0.00	0.00	-0.45	0.00	0.00	-0.45	0.00	0.00	-0.45
NY	-0.35	-0.35	-0.35	-0.35	-0.35	-0.35	-0.35	-0.35	-0.69
PA	0.00	-0.41	-0.41	0.00	-0.41	-0.41	0.00	-0.41	-0.41
VA	0.00	-0.46	-0.46	0.00	-0.46	-0.46	0.00	-0.46	-0.46
WA	-0.37	-1.10	-1.83	-0.37	-1.10	-2.19	-0.37	-1.10	-2.19
Export	-0.11	-0.27	-0.49	-0.11	-0.27	-0.49	-0.11	-0.27	-0.49
Import	0.19	0.16	0.14	0.19	0.16	0.14	0.19	0.16	0.14
Total demand	**-0.08**	**-0.18**	**-0.31**	**-0.08**	**-0.18**	**-0.31**	**-0.08**	**-0.18**	**-0.31**
Retail Price (Non-harvest)									
Atlanta	1.61	3.95	7.03	1.61	3.95	7.03	1.61	3.95	7.03
Los Angeles	1.60	4.04	7.32	1.60	4.04	7.32	1.60	4.04	7.32
New York	1.60	3.86	6.92	1.60	3.86	6.92	1.60	3.86	6.92
Philadelphia	1.60	3.86	6.92	1.60	3.86	6.92	1.60	3.86	6.92
Retail Price (Harvest)									
Atlanta	0.45	0.38	0.38	0.45	0.38	0.38	0.45	0.38	0.38
Los Angeles	0.71	0.63	0.55	0.71	0.63	0.55	0.71	0.63	0.55
New York	0.45	0.37	0.37	0.45	0.37	0.37	0.45	0.37	0.37
Philadelphia	0.45	0.37	0.37	0.45	0.37	0.37	0.45	0.37	0.37
Total Welfare	**-0.50**	**-1.14**	**-1.98**	**-0.50**	**-1.14**	**-1.98**	**-0.50**	**-1.14**	**-1.98**

Table 1.10 Percent changes (%) in CO2 emissions under a carbon tax scheme

	Scenario I			Scenario II					
	Carbon Tax			Carbon Tax with Sequestration Option (i)			Carbon Tax with Sequestration Option (ii)		
Tax rates	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**
Emissions from Production									
CA	-0.78	-2.22	-4.20	-0.79	-2.27	-4.28	-0.57	-2.12	-4.37
MI	-0.03	-0.08	-0.15	-0.03	-0.08	-0.16	-0.02	-0.09	-0.18
NY	-0.04	-0.07	-0.12	-0.04	-0.07	-0.12	-0.04	-0.08	-0.13
PA	-0.04	-0.10	-0.17	-0.04	-0.10	-0.18	-0.03	-0.10	-0.20
VA	-0.02	-0.09	-0.19	-0.02	-0.10	-0.20	0.00	-0.10	-0.24
WA	-0.02	-0.09	-0.19	-0.02	-0.10	-0.19	-0.01	-0.09	-0.20
Emissions from Regular Storage									
CA	-30.44	-30.44	-30.44	-30.44	-30.44	-30.44	-30.44	-30.44	-30.44
MI	8.98	4.52	7.10	8.98	4.52	7.09	8.99	4.51	7.06
NY	7.16	7.47	7.41	7.16	7.46	7.40	7.16	7.46	7.38
PA	-58.59	-58.74	-62.71	-58.59	-58.74	-62.72	-58.56	-58.76	-62.79
VA	-33.28	-33.72	-34.32	-33.28	-33.75	-34.37	-33.14	-33.78	-34.64
WA	1.16	1.91	2.32	1.16	1.91	2.32	1.16	1.91	2.33
Emission from CA storage									
CA									
MI	-15.07	-7.76	-12.26	-15.07	-7.76	-12.26	-15.07	-7.76	-12.26
NY	-17.96	-18.84	-18.84	-17.96	-18.84	-18.84	-17.96	-18.84	-18.84
PA	37.68	37.68	40.11	37.68	37.68	40.11	37.68	37.68	40.11
VA	6.57	6.57	6.57	6.57	6.57	6.57	6.57	6.57	6.57
WA	0.61	-0.45	-1.28	0.61	-0.45	-1.28	0.61	-0.45	-1.28
Emissions Activity									
Production	-0.06	-0.20	-0.38	-0.06	-0.20	-0.38	-0.04	-0.19	-0.40
Regular Storage	0.05	0.34	0.71	0.05	0.34	0.71	0.05	0.34	0.71
CA Storage	-0.31	-0.93	-1.75	-0.31	-0.93	-1.75	-0.31	-0.93	-1.75
Transportation	0.43	0.24	0.08	0.41	0.21	0.04	0.41	0.25	0.09
Total Emission	**-0.09**	**-0.34**	**-0.68**	**-0.09**	**-0.35**	**-0.69**	**-0.08**	**-0.35**	**-0.71**

Decreases in apple production caused by a newly imposed carbon tax also lead to decreases in total demand. Total demand falls by 0.08, 0.18 and 0.31 percent, in response to a carbon tax of $25, $100 and $200 per mt C, respectively. These changes in total demand are slightly greater than those in total supply. The changes in apple production in each region give rise to different responses in retail markets between the non-harvest and harvest seasons. When a carbon tax of $25, $100 and $200 per mt C is imposed, retail prices during the non-harvest season increase by about 1.6, 4.0 and 7.0 percent, respectively. In contrast, during the harvest season, they increase only by about 0.5, 0.5 and 0.5 percent, respectively. These results indicate that the impacts of a carbon tax seem to affect more those who purchase fresh apples in the non-harvest season than those who purchase fresh apple in the harvest season. These results may be due to changes in apple storage decisions. Recall that CO2 emission rates of apples put into CA storage are higher than those that are put into regular storage. Therefore, a carbon tax increases the relative cost of apples put into CA storage, perhaps inducing less utilization of CA storage. Consequently, retail prices of fresh apples in the non-harvest season may increase more than those in the harvest season if these costs are successfully passed on to consumers.

The overall social costs generated by the introduction of a carbon tax, measured by social welfare losses increase with the carbon tax rate. Our simulation results show that social costs increase by 0.5, 1.14 and 1.98 percent for a carbon tax rate of $25, $100 and $200 per mt C, respectively.

Table 1.10 presents the percent change in CO2 emissions corresponding to changes in production and consumption under a carbon tax scheme. Given that CO2 emissions are proportional to apple production, total CO2 emissions from production activities decrease by 0.06, 0.20 and 0.38 percent under a carbon tax of $25, $100 and $200 per mt C, respectively. As predicted by changes in retail prices, there are changes in the amount of apples kept in regular and CA storage. Total CO2 emissions generated from regular storage increase slightly by 0.05, 0.34 and 0.71 percent for a tax rate of $25, $100 and $200 per mt C, respectively. In contrast, CO2 emissions from CA storage decrease by 0.31, 0.93 and 1.75 percent for the same tax rates, respectively. These results imply that storage managers seem to avoid increases in CA storage costs caused by a carbon tax by selling more apples in the harvest season. These lead to greater increases in retail prices for fresh apples in the non-harvest season than in the harvest season.

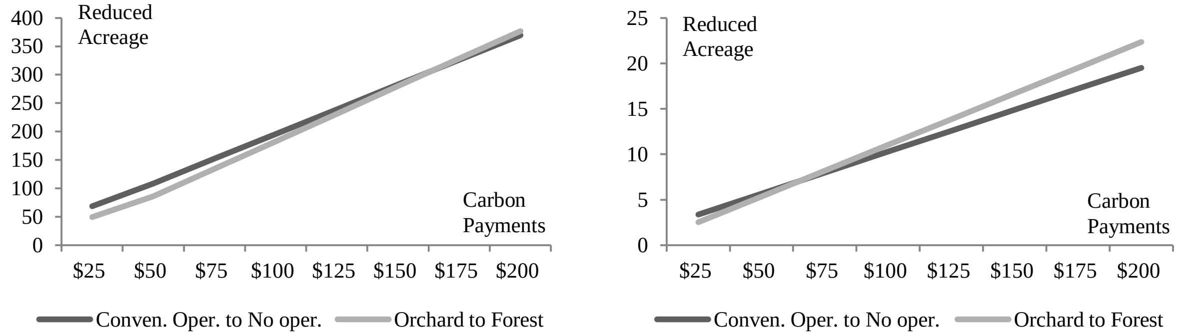
The results also show substantial differences across supply regions regarding CO2 emissions from storage activities. For instance, in Michigan, New York and Washington, the three largest apple producer states, CO2 emissions from regular storage increase, while emissions from CA storage decrease. That is, these three states seek to decrease some of the storage costs caused by a carbon tax by increasing fresh apples sale in the harvest season. In contrast, in Pennsylvania and Virginia, CO2 emissions from regular storage decrease, but those from CA storage increase. That is, these two states rather increase the amount of fresh apples in CA storage in response to increases in retail prices in the non-harvest season.

Furthermore, the simulation results suggest that changes in storage behaviors seem to affect distribution routes from supply regions to consumption locations. Table 1.10 shows that even though total supply and total demand fall, total emissions from transportation activities rather increase by 0.43, 0.24 and 0.08 percent for a carbon tax of $25, $ 100 and $200 per mt C, respectively. This suggests that the inter-temporal reallocation of apple storage causes the Food Miles to increase. Overall, the results from Scenario I suggest that the introduction of a carbon tax of $25, $ 100 and $200 per mt C leads to decreases in fresh apple supply-chain-wide emissions by 0.09, 0.34 and 0.68 percent, respectively.

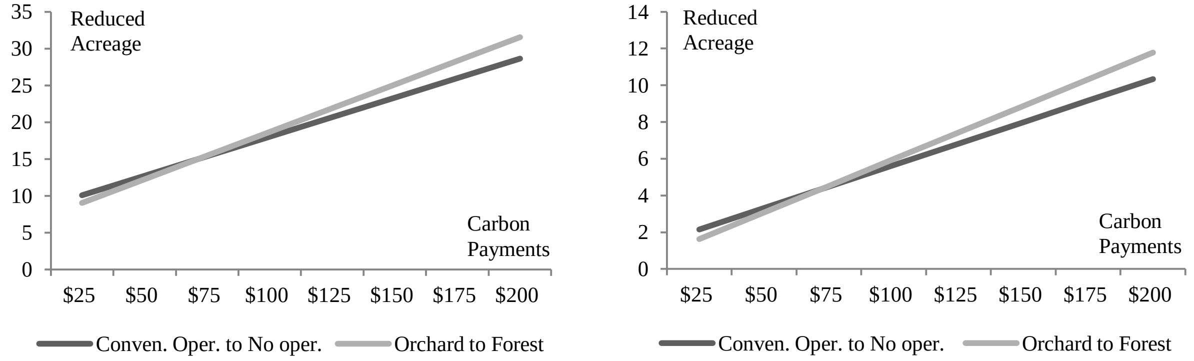
Table 1.9 and 1.10 also present empirical results from Scenario II where two alternative carbon sequestration incentives (i.e. option (i) refers to no orchard operation and option (ii) represents changes in land use from orchard to forest) are incorporated into a carbon tax scheme. In this scenario, if benefits from sequestering CO2 in the soil or biomass through decreases in apple production are greater than the income generated by apple production, orchard owners may participate in a carbon sequestration incentive program. The results show that the overall effects of CO2 sequestration on apple production and the corresponding emission reductions are modest relative to Scenario I (i.e. carbon tax scheme where carbon sequestration options are not considered). Specifically, at a carbon tax and carbon price level of $200 per mt C, sequestration options (i) and (ii) lead to additional reductions in apple production of 0.01 and 0.02 percent, respectively, relative to a carbon tax scheme without carbon sequestration. At carbon price of $200 per mt C, total emissions are reduced 0.01 and 0.03 percent more, relative to Scenario I. These results suggest that carbon payment levels ranging from $25 to $200 per mt C may not provide enough incentive for orchard owners to participate in the carbon sequestration program.

Regardless of the modest impacts of carbon sequestration in these payment ranges, the results of Scenario II provide important information about regional differences of the economic impacts between two carbon sequestration options. Recall that carbon sequestration rates of option (ii) are greater than those of option (i). In contrast, option (ii) requires installation costs for afforestation. In this regard, the scope to which each supply region would participate in carbon sequestration programs may differ based on carbon sequestration rates and carbon payment levels for sequestrated CO2. Therefore, it is meaningful to find differences in economic scope between alternative carbon sequestration options by each supply region. Figure 1.2 exhibits total acreage participating in both carbon sequestration programs by each payment level.

California Michigan



New York Pennsylvania



Virginia Washington

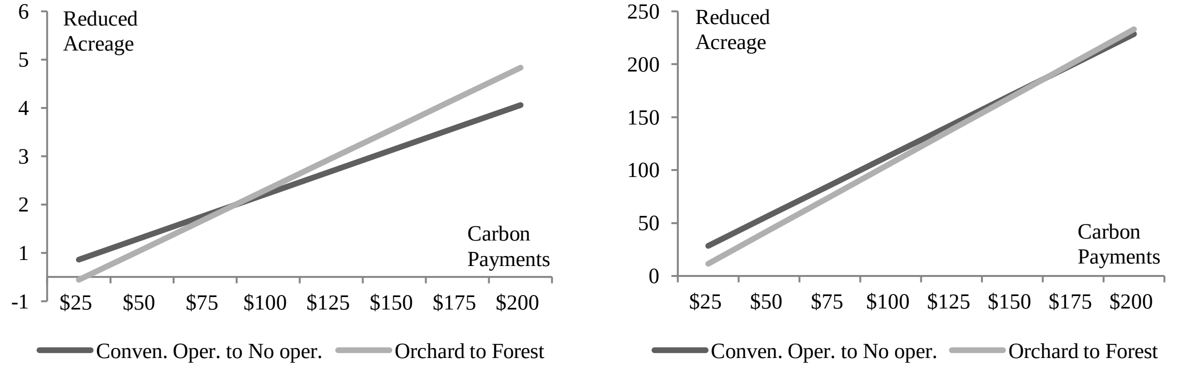


Figure 1.2 Reduced acreages by carbon payment levels between carbon sequestration options

The results suggest that at carbon payment levels between $75 and $100 per mt C, conversion of option (i) to option (ii) seems to be more effective in Michigan, New York, Pennsylvania and Virginia. In contrast, carbon payment levels of around $175 per mt C seem to be a threshold price that induces conversion of option (i) to option (ii) in California and Washington.

1.5.2 Cap-and-Trade

We now examine the economic impacts of a cap-and-trade program targeting CO2 emission reductions in the fresh apple supply chain. As described in Section 4, we consider two possible scenarios regarding a cap-and-trade program: Scenario III examines impacts of a cap-and-trade program and Scenario IV adds carbon sequestration incentive programs to a cap-and-trade system. In both scenarios, CO2 emissions are restricted by a predetermined CO2 emission reduction plan based on CO2 emissions from the baseline solution. Like a carbon tax scheme, a cap-and-trade program is only applied to production and storage segments.

Table 1.11 presents the percent changes in outputs and prices under Scenario III employing various CO2 emission permit prices. The results show that the implementation of a cap-and-trade system leads to decreases in total supply because CO2 emission reductions are primarily accomplished by decreases in apple production. At permit prices of $25, $100 and $200 per mt C, total apple supply under Scenario III decreases by 0.03, 0.13 and 0.27 percent, respectively. These figures are slightly smaller than those under a carbon tax scheme in Scenario I. These results imply that emission permit prices ranging from $25 to $200 per mt C seem to be insufficient to have the same effect as a carbon tax. This is because apple production is still more profitable than selling emission permits in these permit price ranges. Reduced apple production caused by a cap-and-trade program leads to modest decreases in producer prices. Like a carbon tax scheme, California seems to be more responsive to the introduction of a cap-and-trade program.

The impacts on total demand are also modest relative to total demand in the baseline model. However, these effects are about two times less than those under a carbon tax in Scenario I. Specifically, there are seasonal differences in retail prices. Unlike a carbon tax scheme in Scenario I, retail prices in the harvest season for selected consumption locations are unchanged or rather decrease (at permit price of $200 per mt C) under a cap-and-trade system, compared to the baseline. In contrast, retail prices in the non-harvest season rise under a cap-and-trade program. This is similar to retail prices under a carbon tax in Scenario I. But the magnitudes of a rise in retail price in these locations is much smaller than those in Scenario I. These results show different apple storage allocations between regular and CA storage in comparison to a carbon tax scheme in Scenario I. Recall that a carbon tax raises the storage costs. Therefore, storage managers may avoid cost increases by selling more fresh apples in the harvest season under a carbon tax policy. In contrast, under a cap-and-trade system, storage managers may buy or sell their emission permits from or to other practitioners. Therefore, they may adjust their storage plan based on the permit price level.

The social costs caused by the introduction of a cap-and-trade system, measured by welfare losses increase by 0.15, 0.59 and 1.17 percent for permit prices of $25, $100, and $200 per mt C, respectively. These are slightly smaller than those from a carbon tax scheme, mainly due to benefits from trading emission permits between practitioners inside and outside supply chains.

Table 1.11 Percent changes (%) in output and price under a cap-and-trade program

	Scenario III			Scenario IV					
	Cap-and-Trade			Cap-and-Trade with Sequestration Option (i)			Cap-and-Trade with Sequestration Option (ii)		
Permit Prices	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**
Total Supply	**-0.03**	**-0.13**	**-0.27**	**-0.03**	**-0.13**	**-0.27**	**-0.01**	**-0.13**	**-0.28**
Quantity									
CA	-0.33	-1.77	-3.74	-0.34	-1.81	-3.82	-0.12	-1.66	-3.91
MI	-0.02	-0.07	-0.14	-0.02	-0.07	-0.15	-0.01	-0.08	-0.17
NY	-0.01	-0.04	-0.09	-0.01	-0.04	-0.09	-0.01	-0.05	-0.10
PA	-0.02	-0.08	-0.16	-0.02	-0.08	-0.16	-0.01	-0.09	-0.19
VA	-0.02	-0.10	-0.20	-0.02	-0.10	-0.21	0.00	-0.11	-0.25
WA	-0.02	-0.09	-0.19	-0.02	-0.10	-0.19	-0.01	-0.09	-0.20
Producer Price									
CA	-0.81	-3.51	-7.03	-0.81	-3.51	-7.30	-0.27	-3.24	-7.30
MI	0.00	0.00	-0.45	0.00	0.00	-0.45	0.00	0.00	-0.45
NY	-0.35	-0.35	-0.35	-0.35	-0.35	-0.35	-0.35	-0.35	-0.35
PA	0.00	0.00	-0.40	0.00	0.00	-0.40	0.00	0.00	-0.40
VA	0.00	-0.46	-0.46	0.00	-0.46	-0.46	0.00	-0.46	-0.46
WA	-0.36	-1.09	-1.82	-0.36	-1.09	-2.19	-0.36	-1.09	-2.19
Export	-0.03	-0.11	-0.23	-0.03	-0.11	-0.23	-0.03	-0.11	-0.23
Import	-0.01	-0.02	-0.04	-0.01	-0.02	-0.04	-0.01	-0.02	-0.04
Total Demand	**-0.02**	**-0.07**	**-0.13**	**-0.02**	**-0.07**	**-0.13**	**-0.02**	**-0.07**	**-0.13**
Retail Price (Non-harvest)									
Atlanta	0.44	1.68	3.29	0.44	1.68	3.29	0.44	1.68	3.29
Los Angeles	0.38	1.68	3.43	0.38	1.68	3.43	0.38	1.68	3.43
New York	0.36	1.60	3.20	0.36	1.60	3.20	0.36	1.60	3.20
Philadelphia	0.44	1.60	3.28	0.44	1.60	3.28	0.44	1.60	3.28
Retail Price (Harvest)									
Atlanta	0.00	0.00	-0.08	0.00	0.00	-0.08	0.00	0.00	-0.08
Los Angeles	-0.08	-0.08	-0.16	-0.08	-0.08	-0.16	-0.08	-0.08	-0.16
New York	0.00	0.00	-0.07	0.00	0.00	-0.07	0.00	0.00	-0.07
Philadelphia	0.00	0.00	-0.07	0.00	0.00	-0.07	0.00	0.00	-0.07
Total Welfare	**-0.15**	**-0.59**	**-1.17**	**-0.15**	**-0.59**	**-1.17**	**-0.15**	**-0.59**	**-1.17**

Table 1.12 Percent change (%) in CO2 emissions under a cap-and-trade program

	Scenario III			Scenario IV					
	Cap-and-Trade			Cap-and-Trade with Sequestration Option (i)			Cap-and-Trade with Sequestration Option (ii)		
Permit Prices	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**	**$25**	**$100**	**$200**
Emissions from Production									
CA	-0.33	-1.77	-3.74	-0.34	-1.81	-3.82	-0.12	-1.66	-3.91
MI	-0.02	-0.07	-0.14	-0.02	-0.07	-0.15	-0.01	-0.08	-0.17
NY	-0.01	-0.04	-0.09	-0.01	-0.04	-0.09	-0.01	-0.05	-0.10
PA	-0.02	-0.08	-0.16	-0.02	-0.08	-0.16	-0.01	-0.09	-0.19
VA	-0.03	-0.10	-0.20	-0.03	-0.10	-0.21	0.00	-0.11	-0.25
WA	-0.02	-0.09	-0.19	-0.02	-0.10	-0.19	-0.01	-0.09	-0.20
Emissions from Regular Storage									
CA	-0.37	-1.46	-2.92	-0.37	-1.49	-2.98	-0.13	-1.38	-3.05
MI	0.10	8.08	1.84	0.10	8.07	1.83	0.11	8.06	1.79
NY	0.02	-2.73	-2.58	0.02	-2.73	-2.58	0.03	-2.74	-2.60
PA	0.39	1.58	3.16	0.39	1.57	3.14	0.42	1.56	3.08
VA	-0.15	-0.61	-34.37	-0.16	-0.63	-34.42	-0.01	-0.66	-34.68
WA	0.07	0.29	1.27	0.07	0.29	1.28	0.05	0.28	1.29
Emission from CA storage									
CA									
MI	-0.22	-13.67	-3.45	-0.22	-13.67	-3.45	-0.22	-13.67	-3.45
NY	-0.10	6.65	6.12	-0.10	6.65	6.12	-0.10	6.65	6.12
PA	-0.29	-1.15	-2.29	-0.29	-1.15	-2.29	-0.29	-1.15	-2.29
VA	0.00	0.00	6.57	0.00	0.00	6.57	0.00	0.00	6.57
WA	-0.10	-0.38	-1.46	-0.10	-0.38	-1.46	-0.10	-0.38	-1.46
Emissions from Activity									
Production	-0.04	-0.17	-0.35	-0.04	-0.17	-0.36	-0.02	-0.16	-0.37
Regular Storage	0.05	0.20	0.41	0.05	0.20	0.41	0.05	0.20	0.41
CA Storage	-0.11	-0.44	-0.88	-0.11	-0.44	-0.88	-0.11	-0.44	-0.88
Transportation	-0.01	-0.06	-0.26	0.00	-0.18	-0.26	-0.05	-0.09	-0.27
Total Emission	**-0.05**	**-0.21**	**-0.44**	**-0.05**	**-0.23**	**-0.45**	**-0.05**	**-0.21**	**-0.45**

Table 1.12 reports changes in CO2 emissions under Scenario III. Total emissions under a cap-and-trade system decline by 0.05, 0.21 and 0.44 percent for emission permit prices of $25, $100 and $200 per mt C, respectively. The extent to which a cap-and-trade program influences CO2 emission reductions is smaller compared to a carbon tax scheme. These results suggest that emission permit prices ranging from $25 to $200 per mt C are not sufficient to induce production and storage practitioners to participate in an emission trading system.

Changes in CO2 emissions in each segment indicate that a cap-and-trade program affects production regions in a different ways. CO2 emissions from each supply region are all decrease, indicating that all supply regions act as net seller their emission permits relative to the baseline counterpart. They may sell their permits to other segments within supply chain or other industries outside supply chain. However, compared to the results from a carbon tax in Scenario I, some supply regions (California, Michigan, New York and Pennsylvania) discharge more CO2 emissions. This indicates that under a cap-and-trade system, they produce more apples than under a carbon tax scheme.

Similar to a carbon tax scheme, the introduction of a cap-and-trade program also leads to differences in storage patterns. That is, total CO2 emissions from regular storage increase slightly, whereas those from CA storage tend to decrease. However, the simulation results show that, unlike other supply regions, there is differences in storage pattern between a carbon tax and a cap-and-trade in New York. At the permit price level range from $100 and $200 per mt C, CO2 emissions from regular storage decrease, whereas those from CA storage increase in New York. These results are contrary to those in a carbon tax scheme in Scenario I. Moreover, the introduction of a cap-and-trade program does not increase CO2 emissions from transportation activities, indicating that total distances apples move decrease under this policy scheme, unlike a carbon tax scheme in Scenario I.

Table 1.11 also presents the simulation results of scenario IV where two carbon sequestration options are incorporated into a cap-and-trade program. An incentive for carbon sequestration to orchard owners may promote further reductions in apple production. The results show that for carbon payments ranging from \$25 to \$200 per mt C, a carbon sequestration incentive program leads to modest decreases in apple production, although with slight differences across supply regions. Similar to the results of Scenario II, at low carbon payment levels, carbon sequestration option (i) has a greater effect on farm-level production, inducing more participation of orchard owners. In contrast, at higher carbon payment levels, carbon sequestration option (ii) seems to be more efficient due to the installation costs for preparing afforestation. However, similar to Scenario IV, the overall impacts of the carbon payments incentive on outputs, prices and welfare level are modest for carbon payment levels between \$25 and \$200 per mt C. This indicates that carbon payment levels in this range for carbon sequestration do not encourage apple production practitioners to participate in the carbon sequestration incentive program.

Table 1.12 also reveals percent changes in CO2 emissions under Scenario IV. The extent of decreases in CO2 emissions by the introduction of Scenario IV are very similar to the results from Scenario III. The notable differences between sequestration options are only observed at farm-level production. The results show that at carbon payments between \$25 and \$100 per mt C, sequestration option (i) is more effective, whereas at carbon payments of \$200 per mt C, sequestration option (ii) seems to encourage more participation in a carbon sequestration program. (Figure 3.3).

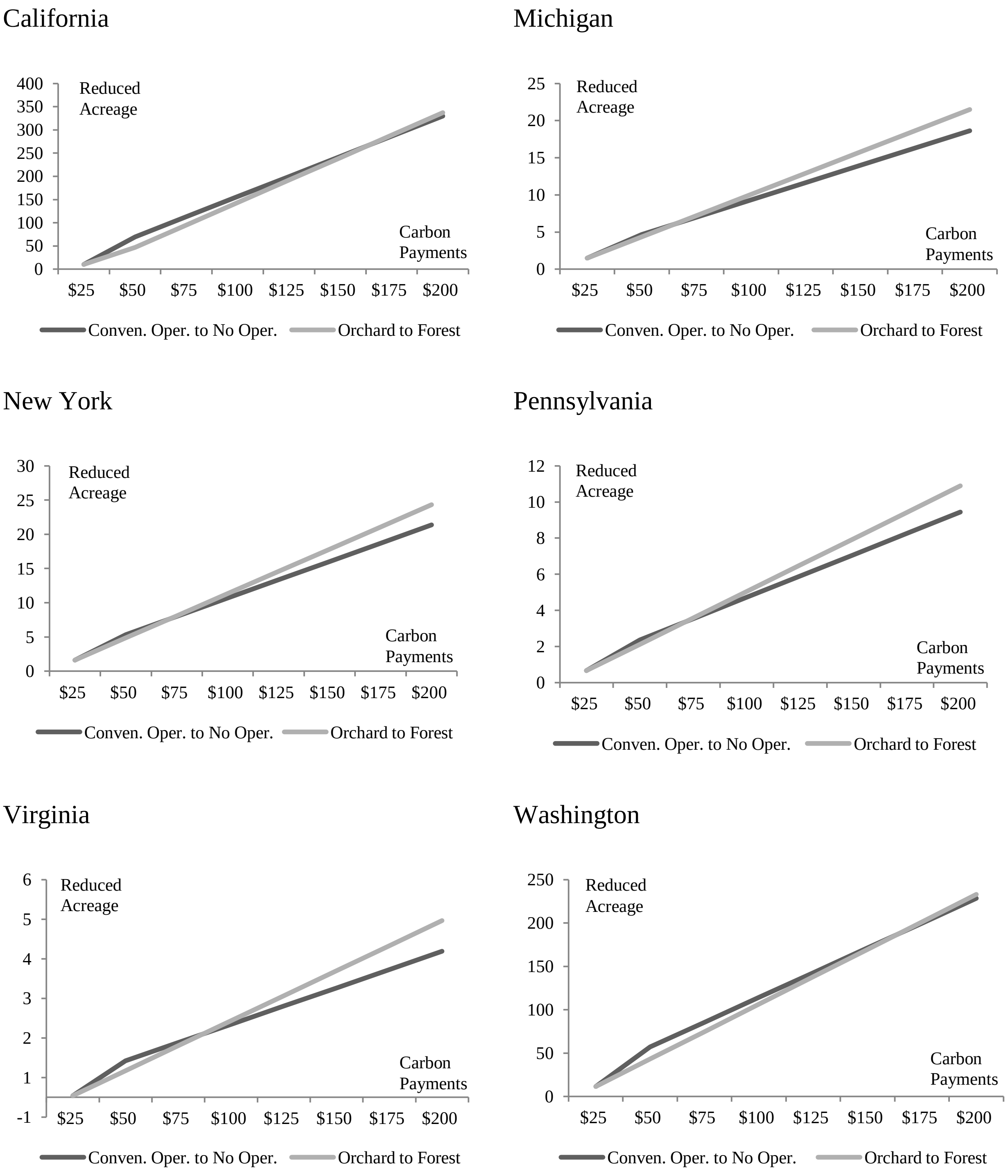


Figure 3.3. Reduced acreages by carbon payment levels between carbon sequestration options under a cap-and-trade program

1.5.3 Cost effectiveness

In this section, we compare the trade-off relationship between CO2 emissions reduction and social costs by examining the degree of cost effectiveness in each policy scenario. Figure 1.4 exhibits the costs per CO2 emissions reduction (CO2 gram) under each scenario.

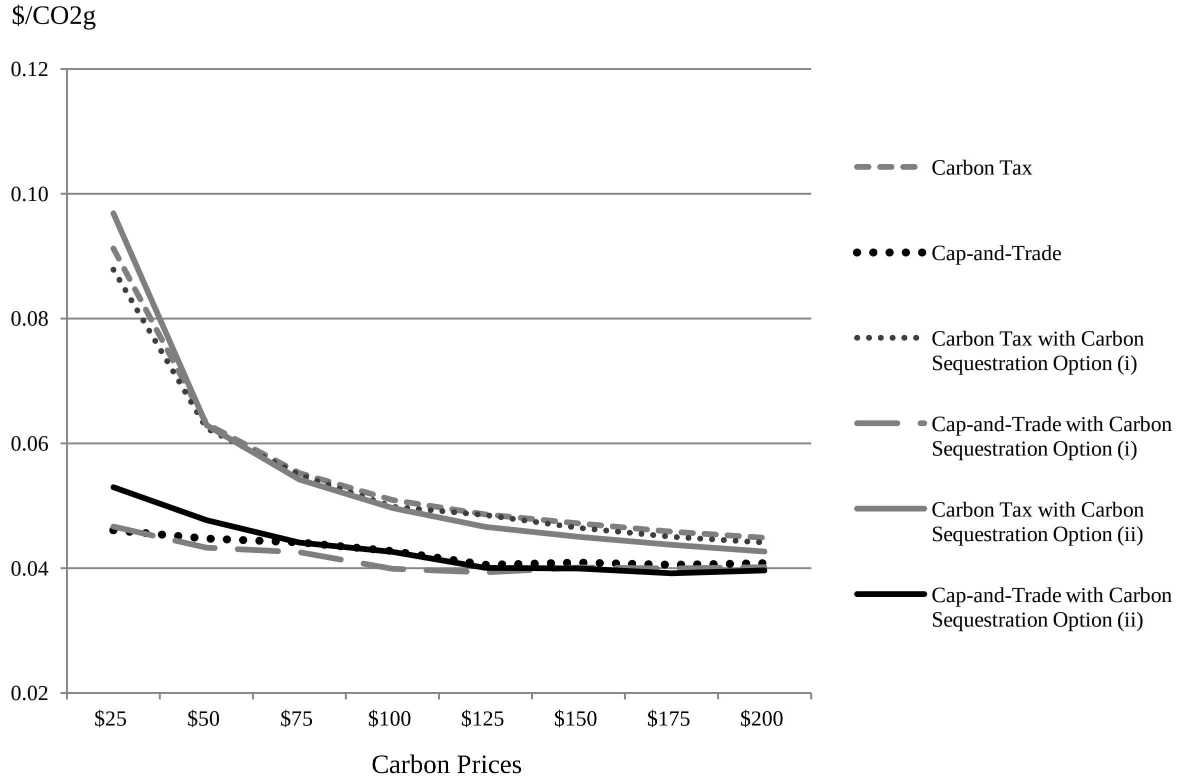


Figure 1.4 Costs per CO2 emissions reduction (g) under carbon prices levels and policy scenarios

These results indicate that, in general, a cap-and-trade system with or without carbon sequestration incentive options in more cost-effective in reducing CO2 emissions from all activities in the fresh apple supply chain, for all carbon prices ranges examined. A cap-and-trade program excluding carbon sequestration incentives seems to be the most cost-effective policy intervention at carbon price of $25 per mt C among all policies. However, if carbon price ranges between $50 and $150 per mt C, a cap-and-trade program with carbon sequestration option (i) is most efficient in mitigating CO2 emissions. On the other hand, for carbon price range between $150 and $200 per mt C, a cap-and-trade system including carbon sequestration option (ii) seems to require the smallest social cost to mitigate a unit amount CO2.

These differences between carbon sequestration options in terms of cost efficiency are also observed in a carbon tax scheme. However, the overall cost efficiencies are smaller than those under a cap-and-trade program. The results show that a carbon tax scheme with sequestration option (ii) leads to the largest social costs to reduce CO2 emissions at a carbon price of $25 per mt C. But, as a carbon price increases, this option seems to be most cost effective than option (i) in terms of CO2 emission reduction among carbon tax policies. In addition, our results imply that there are differences in cost efficiency patterns between a carbon tax and a cap-and-trade system in terms of sequestration options. That is, under a cap-and-trade program, only at a relatively higher payment level (e.g. $150-$200), sequestration option (ii) is cost-effective. However, under a carbon tax scheme, at even lower carbon payment levels (e.g. $50-$200), this option seems to be acceptable compared to sequestration option (i).

1.6 Conclusions

Previous studies to measure CO2 emissions generated by activities along food supply chains such as Food-Miles and LCA provide piecemeal information regarding the tradeoffs relationship between economic performance and environmental impacts. To overcome this problem, we built an optimization model of the U.S. fresh apple supply chain employing a spatial equilibrium framework. We examined the impacts of alternative CO2 emission policies, including a carbon tax scheme and a cap-and-trade program, on fresh apple supply chains in terms of the extent of CO2 emission reduction contributions and cost effectiveness to reduce CO2 emissions. We also investigated the potential impact of carbon sequestration incentives potentially involved in agricultural production practices.

The results show that all CO2 emission policies examined in this study lead to only slight decreases in total CO2 emissions from the fresh apple supply chain. These results are brought about primarily by decreases in apple production in each supply region. Given regional differences in price elasticities of supply, production costs and sequestration rates, California seems to be most affected by the introduction of CO2 emission policies. Our results suggest that a carbon tax scheme may lead to more decreases in apple production than a cap-and-trade program at the same carbon prices. These results indicate that, compared to a carbon tax scheme, benefits from selling emission permits in some regions are still less than those from producing more apples by purchasing additional emission permits from other entities or other sectors. These results are obtained when carbon prices range between $25 and $200 per mt C.

Furthermore, impacts of CO2 emission polices seem to be passed on primarily to consumers who purchase fresh apples in the non-harvest season. This is mainly due to changes in apple storage usage because CO2 emission policies lead to greater utilization of short-term storage and lesser use of CA storage. However, as in apple production, the extent of impacts on fresh apple demand is larger in a carbon tax scheme than in a cap-and-trade program. In addition, carbon sequestration options targeting CO2 emission reductions appear to decrease total CO2 emissions a bit more compared to a carbon tax or a cap-and-trade system without carbon sequestration options. But, the magnitude of CO2 emission reductions is modest under carbon payments ranging from $25 to $200, regardless of a carbon tax or a cap-and-trade. Our results suggest that a cap-and-trade system is more cost-efficient than a carbon tax scheme for CO2 emission mitigation in the fresh apple supply chain, regardless of the presence of CO2 sequestration options.

The models and simulation results in this study provide valuable insights for decision makers promoting industry- or sector-specific initiatives intended to reduce their carbon footprint. Furthermore, the model developed here can assist policy-makers who are interested in the identification of socially acceptable carbon prices and tax levels satisfying both CO2 emission reductions and social welfare losses through the simulation of various policy scenarios. But there are limitations that should be addressed in future research. First, our study should be extended to a general equilibrium framework to take into account substitutability between fruit commodities, revenue recycling effects from tax revenues or emission permit sales and policy interaction effects between other sectors. Second, other emission permit allocation rules such as lump-sum grandfathering, output based allocation, value-added based allocation and historical emissions based allocation should be also examined because each has different potential regarding CO2 emission reductions and regional or sectoral competitiveness.

Appendix 1.1 Spatial price equilibrium model

For a better understanding of a spatial price equilibrium framework, we present a simple model with a single commodity and two separated locations. Consider a perfectly competitive industry with two representative firms. Two firms are assumed to be located in two regions. These firms are price takers in this commodity market. Assume that an inverse demand (supply) function in each region *i* is linear as follows:

$$p_i^{d(s)} = p_i^{d(s)}(q_i^{d(s)}), \tag{A1.1}$$

where $p_i^{d(s)}$ is demand (supply) prices and $q_i^{d(s)}$ is demand (supply) quantities in each region *i*. Let $c(x)$ be the vector of per unit cost with a constant marginal cost structure

in all practices. To simplify, this cost is transportation cost from region *i* to region *j*. Then net social welfare is given by:

$$NSW = \sum_i \left[\int_0^{q_i^d} p_i^d(q_i^d)dq_i^d - \int_0^{q_i^s} p_i^s(q_i^s)dq_i^s \right] - \sum_i \sum_j c_{i,j}(x_{i,j}) \cdot x_{i,j}, \tag{A1.2}$$

where $x_{i,j}$ is the quantities shipped from supply region *i* to demand point *j*.

The spatial price equilibrium model solves this quadratic form in equation (A1.2) to maximize social surplus measured as the sum of consumers and producers surplus minus transportation costs. The model yields optimal quantities for supply, demand; inter-regional shipments; producer prices, consumer prices; and social welfare. The objective function in equation (A1.2) is constrained to ensure the balances between supply and demand. First, the quantities produced in supply region *i* must be greater than or equal to total outflows from each supply region *i* to all consumption locations *j*. Second, the quantities consumed in demand location *j* must be less than or equal to total inflows from all supply regions *i* to each demand location *j*. A social planner chooses supply, demand and shipment quantities to maximize social surplus subject to the balance between supply and demand as follows:

$$\underset{q_i^s, q_i^d, x_{i,j}}{Max} \sum_i \left[\int_0^{q_i^d} p_i^d(q_i^d) dq_i^d - \int_0^{q_i^s} p_i^s(q_i^s) dq_i^s \right] - \sum_i \sum_j c_{i,j}(x_{i,j}) \cdot x_{i,j} \tag{A1.3}$$

$$s.t. \quad q_i^d - \sum_i x_{i,j} \le 0 \quad \forall j \, , \tag{A1.4}$$

$$\sum_j x_{i,j} - q_i^s \le 0 \quad \forall i \, . \tag{A1.5}$$

Let λ_1 and λ_2 be the *Lagrangian* multipliers which are associated with equations (A1.4) and (A1.5), respectively. The first order condition of this maximization problem implies that at optimum, the *Lagrangian* multiplier λ_1 and λ_2 represent demand shadow price, p_i^d and supply shadow price, p_i^s, respectively. The transportation costs are determined by the gap between consumer and producer prices at optimum.

Appendix 1.2 The spatial price equilibrium model with policy application

A1.2.1. A carbon tax scheme

Let τ be the carbon tax rate levied on a metric ton of CO2 (mt C) emitted from apple production and storage activities. The newly imposed tax increases the production and storage costs in proportion to tax rates, τ and CO2 emission rates, $\mu_{p.i}$ and $\mu^t_{s,i}$, respectively. That is, the new production and storage costs are $p\hat{c}_{k,i} = pc_{k,i} + \tau \cdot \mu_{p,i}$ and $s\hat{c}^t_{s,i} = sc^t_{s,i} + \tau \cdot \mu^t_{s,i}$, respectively. Therefore, the objective function in equation (1.1) is modified as follows:

$$
\begin{aligned}
Max \quad & \sum_t \sum_k \sum_j \delta^{t-1} \int_0^{qd^t_{k,j}} pd^t_{k,j}(qd^t_{k,j}) dqd^t_{k,j} - \sum_k \sum_i \int_0^{qs_{k,i}} ps_{k,i}(qs_{k,i}) dqs_{k,i} \\
& + \sum_t \delta^{t-1} \int_0^{qim^t} ppim^t(qim^t) dqim^t - \sum_t \delta^{t-1} \int_0^{qex^t} ppex^t(qex^t) dqex^t \\
& - \sum_k \sum_i (pc_{k,i} + \tau \cdot \mu_{p,i}) \cdot qs_{k,i} - \sum_t \sum_s \sum_i \delta^{t-1} \left[(sc^t_{s,i} + \tau \cdot \mu^t_{s,i}) \cdot \left\{ \sum_k sa^t_{k,s,i} \right\} \right] \\
& - \sum_t \sum_i \sum_j \delta^{t-1} \left[tc^t_{i,j} \cdot \left\{ \sum_k \sum_s tfa^t_{k,s,i,j} \right\} \right] - \sum_t \sum_m \sum_j \delta^{t-1} \left[tc^t_{m,j} \cdot \left\{ \sum_k tfam^t_{k,m,j} \right\} \right]
\end{aligned}
\quad , \ \text{(A1.6)}
$$

The new objective function in equation (A1.6) is maximized subject to the same set of constraints as equations (1.2)-(1.7).

A1.2.2. A cap-and-trade program

Production and storage sectors in the fresh apple supply chains are assumed to be regulated under a cap-and-trade program. In this program, aggregate CO2 emissions generated by production and storage activities cannot exceed a predetermined emission cap, $\overline{E} = \sum_i E_i$, where E_i is CO2 emissions emitted from supply region *i*. Let A_i be the lump-sum allocation of CO2 emission permits to the supply region i . Total emission permits distributed to each entity are determined by predetermined CO2 emission reduction plan. That is, $\sum_i A_i = (1-\psi)\overline{E}$, where ψ is the predetermined emission reduction plan. In general, emission permits can be distributed to each entity by grandfathering or auctioning. Under grandfathering, each entity receive emission permits freely based on the allocation rule (e.g. output based or performance based). In contrast, under auctioning, each entity must purchase emission permits to the extent to which they are going to produce apples in emission permit market. We consider in this study, auctioning as an allocation method. Let ω refer to the permit price. Under a cap-and-trade program, the new objective function is given as:

$$\begin{aligned}
Max \quad & \sum_t \sum_k \sum_j \delta^{t-1} \int_0^{qd^t_{k,j}} pd^t_{k,j}(qd^t_{k,j})dqd^t_{k,j} - \sum_k \sum_i \int_0^{qs_{k,i}} ps_{k,i}(qs_{k,i})dqs_{k,i} \\
& + \sum_t \delta^{t-1} \int_0^{qim^t} ppim^t(qim^t)dqim^t - \sum_t \delta^{t-1} \int_0^{qex^t} ppex^t(qex^t)dqex^t \\
& - \sum_k \sum_i pc_{k,i} \cdot qs_{k,i} - \sum_t \sum_s \sum_i \delta^{t-1} \left[sc^t_{s,i} \cdot \left\{ \sum_k sa^t_{k,s,i} \right\} \right] \qquad , \text{(A1.7)} \\
& - \sum_t \sum_i \sum_j \delta^{t-1} \left[tc^t_{i,j} \cdot \left\{ \sum_k \sum_s tfa^t_{k,s,i,j} \right\} \right] - \sum_t \sum_m \sum_j \delta^{t-1} \left[tc^t_{m,j} \cdot \left\{ \sum_k tfam^t_{k,m,j} \right\} \right] \\
& + \sum_i \omega \left[A_i - \mu_{p,i} \sum_k qs_{k,i} - \mu^t_{s,i} \sum_t \sum_k sa^t_{k,s,i} \right]
\end{aligned}$$

where the term, $\sum_i \omega \cdot [A_i - \mu_{p,i} \sum_k qs_{k,i} - \mu_{s,i}^t \sum_t \sum_k sa_{k,s,i}^t]$ describes the benefits from selling or buying emission permits to or from other regions and segments of fresh apple supply chain. In addition, total CO2 emissions from production and storage activities must be less than or equal to the predetermined emission reduction plan. New constraint is added to the set of constraints in equations (1.2)-(1.7).

$$\sum_i [\mu_{p,i} \sum_i qs_{k,i} + \mu_{s,i}^t \sum_t \sum_k sa_{k,s,i}^t] \leq (1-\psi)\overline{E}\,, \tag{A1.8}$$

A1.2.3. Carbon sequestration incentive program

An introduction of carbon sequestration incentive program provides benefits for orchard owners if they participate in the incentive program. Two activities are incorporated into our model: (i) changes in production practice from conventional orchard operation to no orchard operation; (ii) changes in land use from orchard to forest.

Let $\pi_{1,i}$ and $\pi_{2,i}$ represent the sequestration rates for changes in production practice from conventional operation to no operation and for changes in land use from orchard to forest in supply region *i*, respectively. Denote pr by the price of sequestered unit CO2. Under a carbon tax or a cap-and-trade program, if apple production is reduced, orchard owners have a choice between letting no production acres be set-aside land and shifting no production acres to forest.

If orchard owners choose option (i), they receive carbon payments from the regulator for the amounts of sequestered CO2, obtained by activity of option (i) as follows:

$$pr \cdot \left[\sum_i \pi_{1,i} \cdot \left\{ (\overline{qs}_i - \sum_k qs_{k,i}) / aveyield_i \right\} \right], \tag{A1.9}$$

where $\overline{qs}_i$ represents the baseline apple production in region *i*. To implement carbon sequestration option (i), equation (A1.9) can be incorporated into both equation (1.6) under a carbon tax scheme and equation (A1.7) under a cap-and-trade program.

In contrast, if orchard owners participate in option (ii) incentive program, for set-aside lands acquiring from apple production reduction, they convert orchard lands in no production into forest. It requires forest establishment cost. Let $forestc_i$ be the unit cost for forest establishment per acre in region *i*. Then, their benefits are:

$$(pr - forestc_i) \cdot \left[\sum_i \pi_{2,i} \cdot \left\{ (\overline{qs}_i - \sum_k qs_{k,i}) / aveyield_i \right\} \right], \tag{A1.10}$$

Therefore, orchard owners would choose this option, if the price of sequestered CO2 exceeds forest installation costs. Similarly, carbon sequestration option (ii) is implemented by adding equation (A1.10) to equation (1.6) under a carbon tax scheme or equation (A1.7) under a cap-and-trade program.

Appendix 1.3 Estimation results of the LA-ADIS model and cost data tables

Table A1.1 Estimation results of LA-AIDS model for the non-harvest season in the Northeast

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.199	0.011	107.01	<.0001
	η	-0.126	0.002	-78.45	<.0001
	lnp_{gd}	-0.038	0.019	-2.01	0.045
	lnp_{gs}	0.054	0.016	3.32	0.0009
	lnp_{rd}	-0.011	0.009	-1.18	0.2383
	lnp_{gl}	-0.005	0.011	-0.46	0.6448
	lnp_{ot}	-0.001	0.006	-0.11	0.9131
	X/P	0.003	0.002	2.04	0.0414
	MR_{gd}	-0.165	0.001	-114.92	<.0001
	T	0.000	0.003	-0.04	0.9669
GS	*Constant*	1.283	0.034	37.34	<.0001
	η	-0.287	0.009	-33.67	<.0001
	lnp_{gd}	0.054	0.016	3.32	0.0009
	lnp_{gs}	-0.096	0.029	-3.29	0.001
	lnp_{rd}	0.012	0.015	0.8	0.4248
	lnp_{gl}	-0.006	0.015	-0.41	0.6842
	lnp_{ot}	0.036	0.016	2.26	0.0241
	X/P	0.000	0.005	0	0.999
	MR_{gd}	-0.344	0.009	-37.36	<.0001
	T	-0.009	0.009	-1.02	0.3083
RD	*Constant*	1.203	0.007	161.51	<.0001
	η	0.024	0.001	16.41	<.0001
	lnp_{gd}	-0.011	0.009	-1.18	0.2383
	lnp_{gs}	0.012	0.015	0.8	0.4248
	lnp_{rd}	-0.002	0.008	-0.2	0.8436
	lnp_{gl}	-0.002	0.008	-0.2	0.8436
	lnp_{ot}	0.002	0.007	0.28	0.7813
	X/P	-0.007	0.002	-3.39	0.0007
	MR_{gd}	0.332	0.002	193.52	<.0001
	T	0.004	0.004	0.98	0.3264
GL	*Constant*	1.067	0.007	145.44	<.0001
	η	0.013	0.001	10.07	<.0001
	lnp_{gd}	-0.005	0.011	-0.46	0.6448
	lnp_{gs}	-0.006	0.015	-0.41	0.6842
	lnp_{rd}	-0.014	0.011	-1.25	0.2122
	lnp_{gl}	0.032	0.013	2.37	0.0179
	lnp_{ot}	-0.007	0.006	-1.11	0.2651
	X/P	-0.005	0.002	-2.72	0.0066
	MR_{gd}	0.357	0.002	169.99	<.0001
	T	0.004	0.004	1.27	0.2044
# of Observations	1465		R^2	0.93	

Table A1.2 Estimation results of LA-AIDS model for the harvest season in the Northeast

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.080	0.011	94.94	<.0001
	η	-0.110	0.002	-61.64	<.0001
	lnp_{gd}	-0.028	0.016	-1.75	0.0796
	lnp_{gs}	0.040	0.015	2.74	0.0063
	lnp_{rd}	-0.008	0.010	-0.78	0.4347
	lnp_{gl}	0.000	0.011	0.02	0.9806
	lnp_{ot}	-0.005	0.006	-0.82	0.4132
	X/P	-0.004	0.002	-2.19	0.0289
	MR_{gd}	-0.148	0.001	-103.62	<.0001
	T	-0.007	0.004	-1.99	0.0462
GS	*Constant*	1.154	0.030	38.03	<.0001
	η	-0.245	0.008	-32.21	<.0001
	lnp_{gd}	0.040	0.015	2.74	0.0063
	lnp_{gs}	-0.167	0.029	-5.76	<.0001
	lnp_{rd}	0.025	0.018	1.41	0.1595
	lnp_{gl}	0.070	0.018	3.95	<.0001
	lnp_{ot}	0.031	0.014	2.23	0.0261
	X/P	-0.014	0.004	-3.25	0.0012
	MR_{gs}	-0.302	0.008	-37.34	<.0001
	T	-0.007	0.009	-0.82	0.4149
RD	*Constant*	1.072	0.011	95.31	<.0001
	η	0.019	0.002	8.31	<.0001
	lnp_{gd}	-0.008	0.010	-0.78	0.4347
	lnp_{gs}	0.025	0.018	1.41	0.1595
	lnp_{rd}	0.000	0.010	-0.03	0.9723
	lnp_{gl}	0.000	0.010	-0.03	0.9723
	lnp_{ot}	-0.016	0.010	-1.66	0.0969
	X/P	-0.018	0.003	-5.97	<.0001
	MR_{rd}	0.289	0.003	112.44	<.0001
	T	-0.009	0.006	-1.59	0.1131
GL	*Constant*	0.926	0.011	84.93	<.0001
	η	0.012	0.002	5.9	<.0001
	lnp_{gd}	0.000	0.011	0.02	0.9806
	lnp_{gs}	0.070	0.018	3.95	<.0001
	lnp_{rd}	0.007	0.016	0.42	0.6729
	lnp_{gl}	-0.068	0.019	-3.7	0.0002
	lnp_{ot}	-0.009	0.009	-0.98	0.329
	X/P	-0.016	0.003	-5.74	<.0001
	MR_{gl}	0.306	0.003	103.44	<.0001
	T	-0.001	0.006	-0.19	0.8525
# of Observations	1707		R^2		0.83

Table A1.3 Estimation results of LA-AIDS model for the non-harvest season in the Midwest

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.212	0.014	85.73	<.0001
	η	-0.124	0.002	-59.47	<.0001
	lnp_{gd}	-0.074	0.017	-4.28	<.0001
	lnp_{gs}	0.070	0.019	3.76	0.0002
	lnp_{rd}	-0.007	0.010	-0.65	0.5177
	lnp_{gl}	0.005	0.012	0.4	0.6879
	lnp_{ot}	0.006	0.007	0.9	0.3683
	X/P	-0.003	0.002	-1.23	0.2176
	MR_{gd}	-0.167	0.002	-94.32	<.0001
	T	-0.003	0.004	-0.65	0.5137
GS	*Constant*	1.386	0.037	37.46	<.0001
	η	-0.311	0.009	-34.19	<.0001
	lnp_{gd}	0.070	0.019	3.76	0.0002
	lnp_{gs}	-0.212	0.035	-5.98	<.0001
	lnp_{rd}	0.053	0.018	2.89	0.0039
	lnp_{gl}	0.045	0.019	2.4	0.0164
	lnp_{ot}	0.044	0.015	2.85	0.0044
	X/P	-0.010	0.005	-1.98	0.0481
	MR_{gs}	-0.361	0.010	-37.56	<.0001
	T	-0.010	0.009	-1.07	0.285
RD	*Constant*	1.168	0.009	122.91	<.0001
	η	0.025	0.002	12.66	<.0001
	lnp_{gd}	-0.007	0.010	-0.65	0.5177
	lnp_{gs}	0.053	0.018	2.89	0.0039
	lnp_{rd}	-0.020	0.010	-2.09	0.0367
	lnp_{gl}	-0.020	0.010	-2.09	0.0367
	lnp_{ot}	-0.007	0.008	-0.81	0.4175
	X/P	-0.005	0.003	-2.04	0.0417
	MR_{rd}	0.326	0.002	146.36	<.0001
	T	0.010	0.005	1.9	0.0573
GL	*Constant*	1.035	0.010	103.32	<.0001
	η	0.014	0.002	7.95	<.0001
	lnp_{gd}	0.005	0.012	0.4	0.6879
	lnp_{gs}	0.045	0.019	2.4	0.0164
	lnp_{rd}	-0.016	0.013	-1.18	0.2377
	lnp_{gl}	-0.025	0.016	-1.5	0.1344
	lnp_{ot}	-0.009	0.008	-1.17	0.2404
	X/P	-0.005	0.002	-2.25	0.0244
	MR_{gl}	0.346	0.003	119.17	<.0001
	T	0.000	0.005	-0.01	0.992
# of Observations	1113		R^2	0.92	

Table A1.4 Estimation results of LA-AIDS model for the harvest season in the Midwest

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.144	0.013	85.02	<.0001
	η	-0.113	0.002	-50.79	<.0001
	lnp_{gd}	-0.012	0.017	-0.69	0.4881
	lnp_{gs}	0.012	0.018	0.68	0.4987
	lnp_{rd}	0.004	0.011	0.36	0.7226
	lnp_{gl}	-0.005	0.011	-0.44	0.6583
	lnp_{ot}	0.001	0.007	0.13	0.8938
	X/P	-0.012	0.002	-4.92	<.0001
	MR_{gd}	-0.156	0.002	-94.54	<.0001
	T	0.002	0.005	0.36	0.7209
GS	*Constant*	1.062	0.039	27.14	<.0001
	η	-0.226	0.010	-23.5	<.0001
	lnp_{gd}	0.012	0.018	0.68	0.4987
	lnp_{gs}	-0.033	0.033	-0.99	0.3205
	lnp_{rd}	0.001	0.019	0.07	0.944
	lnp_{gl}	0.008	0.019	0.4	0.6865
	lnp_{ot}	0.012	0.016	0.78	0.4342
	X/P	-0.018	0.005	-3.37	0.0008
	MR_{gs}	-0.282	0.010	-27.81	<.0001
	T	0.003	0.010	0.33	0.7436
RD	*Constant*	1.105	0.012	93.16	<.0001
	η	0.023	0.002	9.49	<.0001
	lnp_{gd}	0.004	0.011	0.36	0.7226
	lnp_{gs}	0.001	0.019	0.07	0.944
	lnp_{rd}	-0.005	0.010	-0.56	0.5785
	lnp_{gl}	-0.005	0.010	-0.56	0.5785
	lnp_{ot}	0.006	0.009	0.6	0.5469
	X/P	-0.016	0.003	-5.07	<.0001
	MR_{rd}	0.301	0.003	110.18	<.0001
	T	-0.007	0.006	-1.11	0.2654
GL	*Constant*	0.981	0.012	78.74	<.0001
	η	0.010	0.002	3.99	<.0001
	lnp_{gd}	-0.005	0.011	-0.44	0.6583
	lnp_{gs}	0.008	0.019	0.4	0.6865
	lnp_{rd}	-0.024	0.017	-1.41	0.1575
	lnp_{gl}	0.036	0.017	2.09	0.0372
	lnp_{ot}	-0.014	0.010	-1.51	0.1303
	X/P	-0.025	0.003	-7.56	<.0001
	MR_{gl}	0.317	0.003	91.66	<.0001
	T	-0.005	0.006	-0.74	0.4587
# of Observations	1282		R^2	0.86	

Table A1.5 Estimation results of LA-AIDS model for the non-harvest season in the South

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.232	0.007	169.82	<.0001
	η	-0.130	0.001	-118.42	<.0001
	lnp_{gd}	-0.030	0.009	-3.41	0.0007
	lnp_{gs}	0.026	0.009	2.83	0.0046
	lnp_{rd}	-0.004	0.005	-0.73	0.4641
	lnp_{gl}	0.003	0.007	0.39	0.6967
	lnp_{ot}	0.005	0.005	1.06	0.2878
	X/P	-0.003	0.001	-2.67	0.0077
	MR_{gd}	-0.168	0.001	-180.15	<.0001
	T	0.002	0.002	1.15	0.2513
GS	*Constant*	1.425	0.025	56.79	<.0001
	η	-0.323	0.006	-50.82	<.0001
	lnp_{gd}	0.026	0.009	2.83	0.0046
	lnp_{gs}	-0.066	0.022	-3	0.0027
	lnp_{rd}	0.001	0.010	0.14	0.8851
	lnp_{gl}	-0.006	0.012	-0.46	0.6475
	lnp_{ot}	0.044	0.014	3.12	0.0019
	X/P	-0.010	0.003	-2.89	0.0039
	MR_{gs}	-0.377	0.007	-56.3	<.0001
	T	0.000	0.007	-0.07	0.9455
RD	*Constant*	1.203	0.006	210.97	<.0001
	η	0.024	0.001	20.34	<.0001
	lnp_{gd}	-0.004	0.005	-0.73	0.4641
	lnp_{gs}	0.001	0.010	0.14	0.8851
	lnp_{rd}	0.001	0.005	0.12	0.902
	lnp_{gl}	0.001	0.005	0.12	0.902
	lnp_{ot}	0.001	0.006	0.13	0.8929
	X/P	-0.008	0.002	-5.16	<.0001
	MR_{rd}	0.332	0.001	252.43	<.0001
	T	0.002	0.003	0.81	0.4188
GL	*Constant*	1.062	0.007	158.7	<.0001
	η	0.013	0.001	10.18	<.0001
	lnp_{gd}	0.003	0.007	0.39	0.6967
	lnp_{gs}	-0.006	0.012	-0.46	0.6475
	lnp_{rd}	-0.020	0.009	-2.35	0.019
	lnp_{gl}	0.033	0.012	2.74	0.0062
	lnp_{ot}	-0.010	0.007	-1.38	0.1673
	X/P	-0.009	0.002	-5.58	<.0001
	MR_{gl}	0.353	0.002	182.4	<.0001
	T	-0.004	0.003	-1.26	0.209
# of Observations	2487		R^2		0.93

Table A1.6 Estimation results of LA-AIDS model for the harvest season in the South

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.064	0.009	115.59	<.0001
	η	-0.107	0.002	-68.94	<.0001
	lnp_{gd}	0.001	0.012	0.09	0.9291
	lnp_{gs}	0.016	0.011	1.47	0.1423
	lnp_{rd}	0.001	0.008	0.08	0.9389
	lnp_{gl}	-0.019	0.009	-2.17	0.0303
	lnp_{ot}	0.001	0.006	0.09	0.9301
	X/P	-0.012	0.002	-7.27	<.0001
	MR_{gd}	-0.145	0.001	-126.47	<.0001
	T	0.003	0.003	0.86	0.3877
GS	*Constant*	1.289	0.022	58.02	<.0001
	η	-0.282	0.006	-50.15	<.0001
	lnp_{gd}	0.016	0.011	1.47	0.1423
	lnp_{gs}	-0.081	0.021	-3.93	<.0001
	lnp_{rd}	0.011	0.012	0.94	0.3477
	lnp_{gl}	0.027	0.013	2.14	0.0322
	lnp_{ot}	0.027	0.013	2.06	0.0391
	X/P	-0.017	0.003	-5.11	<.0001
	MR_{gs}	-0.339	0.006	-57.66	<.0001
	T	-0.007	0.007	-0.98	0.329
RD	*Constant*	1.105	0.009	129.45	<.0001
	η	0.021	0.002	11.16	<.0001
	lnp_{gd}	0.001	0.008	0.08	0.9389
	lnp_{gs}	0.011	0.012	0.94	0.3477
	lnp_{rd}	-0.004	0.007	-0.54	0.591
	lnp_{gl}	-0.004	0.007	-0.54	0.591
	lnp_{ot}	-0.004	0.009	-0.46	0.6424
	X/P	-0.020	0.002	-8.47	<.0001
	MR_{rd}	0.301	0.002	156.17	<.0001
	T	0.008	0.005	1.64	0.1018
GL	*Constant*	0.970	0.009	109.97	<.0001
	η	0.010	0.002	5.44	<.0001
	lnp_{gd}	-0.019	0.009	-2.17	0.0303
	lnp_{gs}	0.027	0.013	2.14	0.0322
	lnp_{rd}	-0.021	0.012	-1.79	0.073
	lnp_{gl}	0.031	0.014	2.29	0.0218
	lnp_{ot}	-0.018	0.009	-1.97	0.049
	X/P	-0.021	0.002	-8.99	<.0001
	MR_{gl}	0.318	0.002	136.15	<.0001
	T	0.006	0.005	1.19	0.233
# of Observations	2747		R^2	0.85	

Table A1.7 Estimation results of LA-AIDS model for the non-harvest season in the West

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.267	0.007	188	<.0001
	η	-0.136	0.001	-146.98	<.0001
	lnp_{gd}	-0.052	0.009	-5.9	<.0001
	lnp_{gs}	0.053	0.007	7.16	<.0001
	lnp_{rd}	0.000	0.005	-0.1	0.9224
	lnp_{gl}	-0.004	0.005	-0.75	0.4559
	lnp_{ot}	0.003	0.002	1.55	0.1216
	X/P	0.000	0.001	0.18	0.8557
	MR_{gd}	-0.172	0.001	-194.21	<.0001
	T	-0.001	0.002	-0.4	0.689
GS	*Constant*	1.471	0.030	49.06	<.0001
	η	-0.343	0.008	-45.52	<.0001
	lnp_{gd}	0.053	0.007	7.16	<.0001
	lnp_{gs}	-0.075	0.018	-4.18	<.0001
	lnp_{rd}	-0.020	0.010	-2.08	0.0381
	lnp_{gl}	0.014	0.012	1.11	0.2692
	lnp_{ot}	0.028	0.008	3.5	0.0005
	X/P	-0.011	0.003	-3.72	0.0002
	MR_{gs}	-0.383	0.008	-49.02	<.0001
	T	-0.007	0.006	-1.19	0.2338
RD	*Constant*	1.185	0.007	175.08	<.0001
	η	0.024	0.001	20.94	<.0001
	lnp_{gd}	0.000	0.005	-0.1	0.9224
	lnp_{gs}	-0.020	0.010	-2.08	0.0381
	lnp_{rd}	0.010	0.005	1.89	0.0591
	lnp_{gl}	0.010	0.005	1.89	0.0591
	lnp_{ot}	0.001	0.004	0.25	0.8065
	X/P	-0.003	0.001	-2.17	0.0302
	MR_{rd}	0.327	0.002	204.82	<.0001
	T	0.001	0.003	0.35	0.7258
GL	*Constant*	1.029	0.009	118.15	<.0001
	η	0.009	0.002	5.78	<.0001
	lnp_{gd}	-0.004	0.005	-0.75	0.4559
	lnp_{gs}	0.014	0.012	1.11	0.2692
	lnp_{rd}	-0.012	0.011	-1.04	0.2984
	lnp_{gl}	0.003	0.013	0.24	0.8115
	lnp_{ot}	-0.001	0.005	-0.18	0.8538
	X/P	-0.006	0.002	-3.36	0.0008
	MR_{gl}	0.340	0.002	138.27	<.0001
	T	0.001	0.004	0.33	0.7445
# of Observations	1842		R^2	0.93	

Table A1.8 Estimation results of LA-AIDS model for the harvest season in the West

Budget Share	Variables	Parameter	Std. Err.	t-value	Probability
GD	*Constant*	1.055	0.010	104.6	<.0001
	η	-0.112	0.001	-75.5	<.0001
	lnp_{gd}	0.011	0.012	0.91	0.3625
	lnp_{gs}	0.014	0.010	1.31	0.1888
	lnp_{rd}	-0.016	0.007	-2.23	0.0259
	lnp_{gl}	-0.009	0.006	-1.53	0.1266
	lnp_{ot}	0.000	0.004	-0.06	0.9501
	X/P	-0.004	0.001	-2.74	0.0062
	MR_{gd}	-0.143	0.001	-110.57	<.0001
	T	-0.003	0.003	-1.07	0.2848
GS	*Constant*	1.298	0.027	47.43	<.0001
	η	-0.290	0.007	-41.45	<.0001
	lnp_{gd}	0.014	0.010	1.31	0.1888
	lnp_{gs}	-0.081	0.020	-4.01	<.0001
	lnp_{rd}	0.000	0.012	0.01	0.9927
	lnp_{gl}	0.040	0.012	3.3	0.001
	lnp_{ot}	0.027	0.010	2.71	0.0069
	X/P	-0.012	0.004	-3.18	0.0015
	MR_{gs}	-0.342	0.007	-47.61	<.0001
	T	0.012	0.008	1.51	0.1303
RD	*Constant*	1.032	0.011	96.24	<.0001
	η	0.020	0.002	9.96	<.0001
	lnp_{gd}	-0.016	0.007	-2.23	0.0259
	lnp_{gs}	0.000	0.012	0.01	0.9927
	lnp_{rd}	0.006	0.007	0.93	0.3547
	lnp_{gl}	0.006	0.007	0.93	0.3547
	lnp_{ot}	0.003	0.007	0.44	0.6568
	X/P	-0.016	0.003	-6.37	<.0001
	MR_{rd}	0.280	0.002	113.45	<.0001
	T	-0.003	0.005	-0.63	0.5302
GL	*Constant*	0.963	0.011	88.41	<.0001
	η	0.011	0.002	4.83	<.0001
	lnp_{gd}	-0.009	0.006	-1.53	0.1266
	lnp_{gs}	0.040	0.012	3.3	0.001
	lnp_{rd}	0.001	0.013	0.1	0.9232
	lnp_{gl}	-0.021	0.012	-1.79	0.0742
	lnp_{ot}	-0.010	0.007	-1.44	0.1501
	X/P	-0.023	0.003	-8.68	<.0001
	MR_{gl}	0.312	0.003	112.47	<.0001
	T	-0.006	0.005	-1.18	0.2398
# of Observations	1973		R^2	0.84	

Table A1.9 Average production costs by varieties[a] (Washington)

	Golden Delicious	Granny Smith	Red Delicious	Gala	Fuji (Others)
Variable Costs					
Labor	1,760	1,760	1,570	2,010	1,815
Chemicals	380	380	380	380	380
Operator Labor	285	333	285	333	285
Other	275	275	275	275	275
Total	2,700	2,748	2,510	2,998	2,755
Ownership Costs					
Depreciation	490	490	490	777	775
Interest	770	770	770	1,055	1,055
Taxes & Ins.	150	150	150	210	210
Total	1,410	1,410	1,410	2,042	2,040
Total Cost per Acre	4,110	4,158	3,920	5,040	4,795
Yield (bin)	50	50	40	40	35
$/lb	**0.1025**	**0.1040**	**0.1225**	**0.1575**	**0.1713**

a. Source: Schotzko and Granatstein (2005).

Table A1.10 Unit storage cost by storage facility in each region

	Regular Storage ($/lb)	Controlled Atmosphere Storage ($/lb)
California	0.031	0.044
Michigan	0.019	0.027
New York	0.029	0.042
Pennsylvania	0.020	0.029
Virginia	0.015	0.021
Washington	0.015	0.022

Table A1.11 Transportation costs from supply regions to consumption locations

	ALB	ATL	BAL	BIL	BMH	BOS	BPT	BSE	BUR	CIN
						$/lb				
California	0.056	0.112	0.125	0.057	0.106	0.135	0.130	0.036	0.133	0.106
Michigan	0.072	0.045	0.039	0.063	0.041	0.049	0.044	0.087	0.041	0.027
New York	0.087	0.048	0.023	0.086	0.051	0.027	0.024	0.104	0.024	0.033
Pennsylvania	0.084	0.039	0.014	0.086	0.042	0.029	0.022	0.105	0.030	0.030
Virginia	0.084	0.036	0.016	0.087	0.039	0.033	0.025	0.105	0.033	0.030
Washington	0.064	0.114	0.121	0.041	0.109	0.131	0.125	0.026	0.128	0.102
	CHA	CHE	CHI	CHT	COL	DAL	DEN	DET	FAG	IND
California	0.122	0.058	0.097	0.114	0.120	0.083	0.062	0.107	0.082	0.102
Michigan	0.042	0.057	0.019	0.031	0.045	0.057	0.059	0.018	0.040	0.022
New York	0.038	0.075	0.037	0.032	0.042	0.070	0.077	0.030	0.062	0.035
Pennsylvania	0.029	0.075	0.037	0.025	0.033	0.066	0.076	0.030	0.063	0.032
Virginia	0.026	0.075	0.037	0.022	0.030	0.062	0.076	0.030	0.063	0.032
Washington	0.121	0.056	0.091	0.110	0.122	0.094	0.059	0.103	0.066	0.098
	JAC	KAN	LAG	LTR	LVG	LUS	MCH	MEM	MIA	MIL
California	0.099	0.083	0.028	0.092	0.035	0.104	0.138	0.097	0.137	0.097
Michigan	0.048	0.039	0.100	0.044	0.089	0.026	0.047	0.040	0.072	0.017
New York	0.060	0.055	0.118	0.057	0.107	0.037	0.029	0.052	0.067	0.040
Pennsylvania	0.052	0.052	0.116	0.053	0.107	0.034	0.030	0.048	0.057	0.041
Virginia	0.049	0.052	0.116	0.050	0.107	0.032	0.033	0.044	0.056	0.041
Washington	0.107	0.074	0.054	0.098	0.051	0.101	0.134	0.099	0.141	0.089
	MIN	MOI	NOR	NYC	NWK	OKL	OMH	PHI	PHO	POR
California	0.093	0.083	0.104	0.128	0.127	0.078	0.078	0.126	0.043	0.141
Michigan	0.036	0.032	0.055	0.042	0.041	0.050	0.038	0.040	0.089	0.049
New York	0.053	0.049	0.065	0.023	0.022	0.065	0.055	0.023	0.104	0.032
Pennsylvania	0.053	0.050	0.056	0.021	0.020	0.062	0.055	0.017	0.101	0.033
Virginia	0.054	0.050	0.053	0.023	0.023	0.062	0.056	0.021	0.101	0.036
Washington	0.075	0.079	0.114	0.123	0.123	0.086	0.075	0.122	0.063	0.136
	POT	PRO	RCH	SEA	SLC	STL	SUX	WDC	WMT	
California	0.036	0.136	0.127	0.043	0.040	0.093	0.081	0.124	0.126	
Michigan	0.104	0.045	0.041	0.096	0.075	0.030	0.037	0.038	0.040	
New York	0.122	0.027	0.028	0.119	0.092	0.045	0.060	0.024	0.023	
Pennsylvania	0.122	0.027	0.019	0.119	0.092	0.042	0.060	0.015	0.017	
Virginia	0.123	0.030	0.018	0.119	0.093	0.043	0.060	0.015	0.019	
Washington	0.019	0.132	0.123	0.017	0.040	0.091	0.068	0.120	0.122	

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CHAPTER 2

IMPACTS OF THE END OF THE COFFEE EXPORT QUOTA SYSTEM ON INTERNATIONAL-TO-RETAIL PRICE TRANSMISSION

Abstract

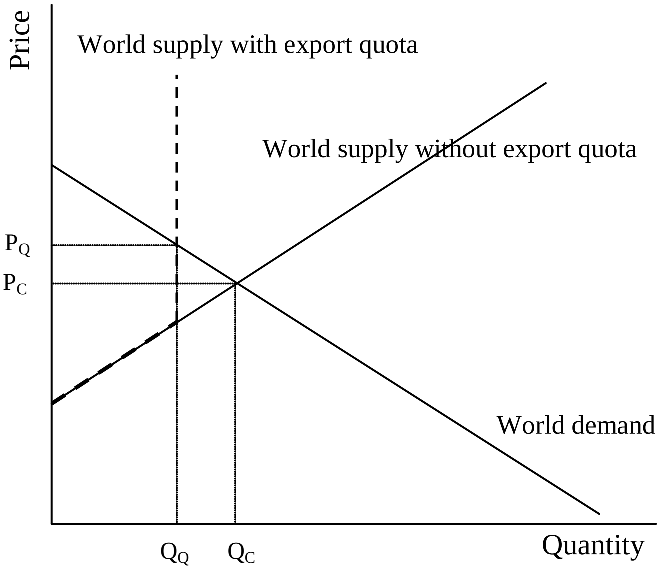
We examine the impact of the end of the coffee export quota system (EQS) on international-to-retail price transmission in France, Germany and the United States, taking into account the existence of long-run threshold effects and short-run price transmission asymmetries (PTAs). We find evidence of threshold effects in two periods (EQS and post-EQS) in the three countries and the presence of short-run PTAs during the post-EQS period in the three countries, but not during the EQS period. Our results indicate that the threshold values are smaller in the post-EQS period, suggesting that retail prices become more responsive to changes in international prices. However, the speed of adjustment toward the long-run equilibrium decreases during the post-EQS period in the three countries. In the short-run, nonlinear impulse response analyses indicate that a shock in international prices was more persistent during the EQS period than in the post-EQS period. Moreover, we find evidence of short-run PTAs in the post-EQS period, with differences across countries. We find support for the "rockets and feathers" principle in the United States; in contrast, retail prices respond faster when international prices are falling in Germany and France. We explain these differences in terms of market structures.

2.1 Introduction

The International Coffee Agreement is a treaty that sets out the objectives and the basic framework for the International Coffee Organization (ICO) member countries, which includes coffee exporting and importing countries. An important policy instrument for the implementation of this agreement was an Export Quota System (EQS), operative from 1962 to 1989. The EQS was introduced to address an oversupply problem in the world coffee market in the 1950s and early 1960s. It was an intergovernmental initiative, supported by importing countries, to stabilize the market and to mitigate the negative political and economic impacts of falling prices for many coffee producing countries in the developing world (ICO 2012).

Figure 2.1 illustrates how the EQS contributed to keep international and retail prices high (Houck 1986). The export quota, Q_Q was set below the competitive equilibrium quantity, Q_C in the international market. The corresponding equilibrium price under the EQS, P_Q was higher than the competitive equilibrium price, P_C. The EQS also led to higher retail prices in importing countries (Panel B of Figure 2.1). Further, Akiyama and Varangis (1990) argue that the EQS contributed to relatively stable international coffee prices, in spite of frequent supply shocks due to weather conditions in the major coffee producing countries. The authors show that the EQS scheme encouraged maintenance of a "buffer stock" to prevent price volatility in the international coffee market.

A. International commodity market



B. Domestic retail market in an importer country

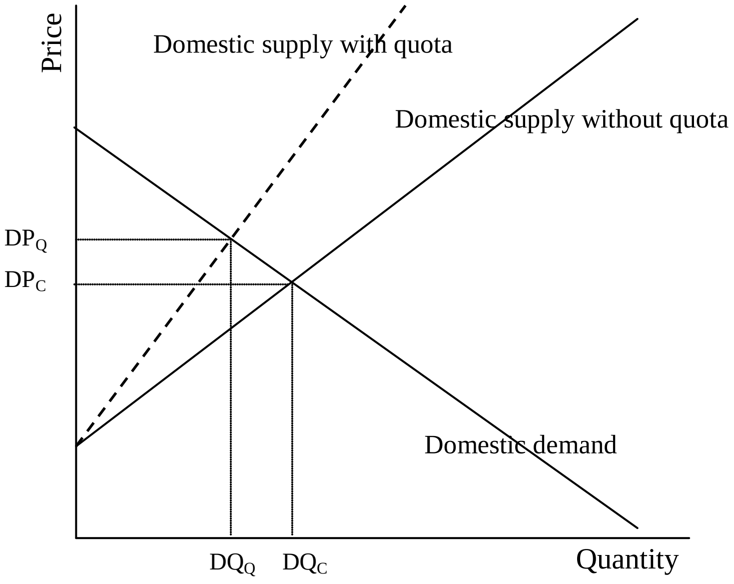


Figure 2.1 The impact of an export quota on the international market and on the domestic market in and importer country

The EQS expired in 1989 due to disagreements on how to allocate the export quota among exporting countries[14] and lack of support from importing countries. Theoretically, termination of the EQS should have resulted in increased export quantity (from Q_Q to Q_C) and reduced international prices (from P_Q to P_C) for coffee beans (Panel A in Figure 2.1). In addition, one would expect that EQS termination should have led to higher price variability in all segments of the supply chain because the market becomes more exposed to exogenous shocks (e.g. climatic conditions). Moreover, EQS elimination should have weakened the ability of exporting countries to set international prices. In contrast, coffee processors and distributors in importing countries should have increased ability to influence the extent to which changes in international prices are passed on to retail prices paid by end consumers. If coffee processors and distributors in importing countries gained market power, it is possible that “rockets and feathers” type of price transmission asymmetries exists in the post-EQS period (Bacon 1991). That is, retail prices rise faster than they fall in response to changes in international prices. Consequently, two relevant empirical questions are: Did price transmission behavior (e.g. the magnitude and the speed) between international and retail prices change after elimination of the EQS? Did price transmission respond differently to increases compared to decreases in international coffee prices?

In this study, we examine whether price transmission between international and retail coffee prices in the three largest coffee importers (France, Germany and the United States) changed in the post-EQS period. Figure 2.2 suggests that the end of the EQS in 1990 may have affected price response in these countries in different ways. For instance, retail prices in the three countries seem to have had a similar relationship with international prices in the EQS period. In

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[14] Under the EQS, the basic export quota of each exporting member was determined each October for a one-year period. The quota was adjusted depending on global market conditions.

contrast, in the post-EQS period, after sharp decreases in international prices in the early 1990s, retail prices in France and the United States decreased accordingly, while retail prices in Germany were more volatile and stayed relatively high.

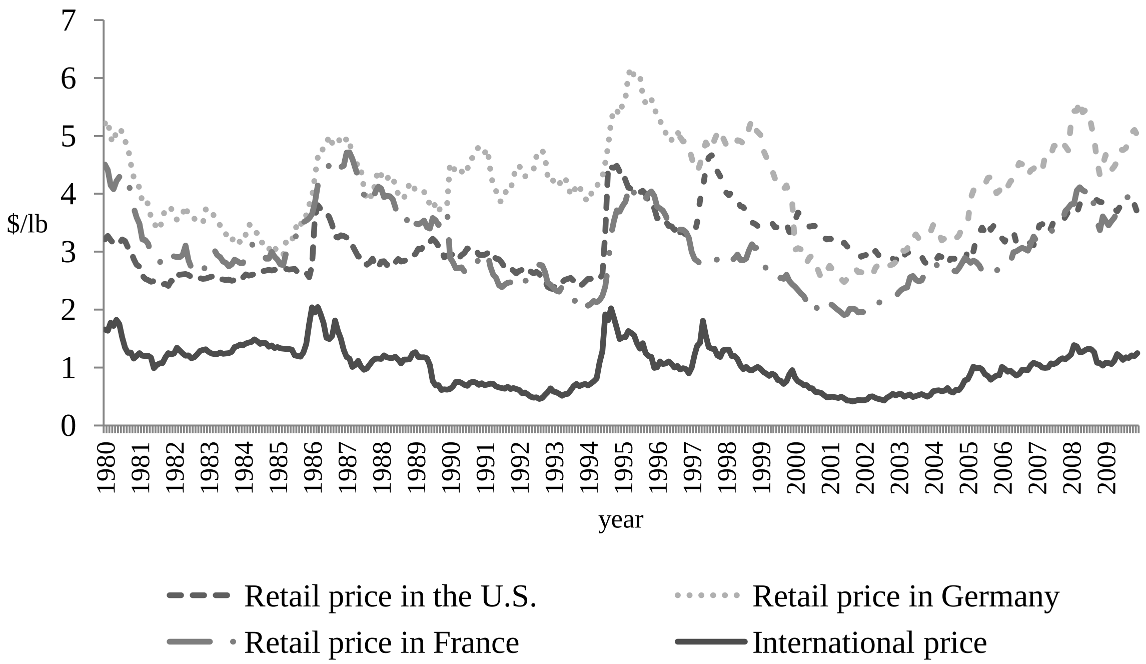

Figure 2.2 Monthly international and retail coffee prices in France, Germany and the United States

There is an extensive literature on price transmission in supply chains for agricultural commodities. Early studies, conducted by Wolffram (1971) and Houck (1977), examined price transmission asymmetries (PTAs) measured by asymmetric supply responses to positive and negative changes in input prices. Observed usually at downstream stages of supply chains, PTAs were found to be closely associated with numerous factors, including market structure, market power, consumer behavior and public policy. Subsequently, von Cramon-Taubadel (1998) pointed out that prices at different segments of the supply chain are often co-integrated and econometric models that ignore this property may generate spurious parameter estimates. The suggested remedy was error correction models (ECMs), which allow for short-run asymmetric price adjustments, to overcome the limitations of Wolffram's and Houck’s approaches (von Cramon-Taubadel and Loy 1996; von Cramon-Taubadel 1998). The standard ECM specification assumes that the dependent variable responds identically to deviations from the long-run price equilibrium regardless of the magnitude of the shock and that adjustments occur in every period (Balke and Fomby 1997). However the presence of any transaction costs between spatially separated markets or other factors generating price friction may result in the existence of thresholds which trigger adjustments toward the long-run equilibrium in response to exogenous shocks (Meyer 2004).

For the purpose here, we focus on two separate dimensions of price transmission in order to understand their relevance for policy impact evaluation. The first dimension posits that price adjustments respond differently to positive and negative exogenous shocks in the short run (i.e. PTAs). The second dimension is that there may be thresholds beyond which price adjustments occur in the long-run (i.e. threshold effects). Our approach involves an error correction model with threshold effects, patterned after work by Tong (1983) and extended by Balke and Fomby (1997). The threshold approach allows us to model price adjustments toward the long-run equilibrium through threshold values estimated using nonparametric methods. In addition, we extend the threshold error correction model by incorporating short-run PTAs.

The overall objective of this study is to examine differences in international-to-retail coffee price transmission behavior during and after the elimination of the EQS in France, Germany and the United States. Specifically, we compare the extent to which shocks in international coffee prices are passed on to retail prices in terms of magnitude and speed, during and after the EQS. In order to understand how price transmission changed with termination of the EQS, we take into account two important features of the price transmission process: the existence of long-run threshold effects and short-run asymmetries. We contribute to the literature by improving our understanding of the impact of policy interventions on price transmission in global supply chains for agricultural commodities.

A number of researchers have examined the impact of the EQS termination at various segments of the coffee supply chain. Akiyama and Varangis (1990) employed simulation methods for the global coffee supply chain to demonstrate that the EQS contributed to the stability of international coffee prices. Krivonos (2004) conducted a co-integration analysis and found that the speed of price transmission between producer and international prices increased in the post-EQS period and the share of retail value going to coffee growers increased after the EQS termination. The analysis also showed that retail prices adjusted faster in response to shocks in international prices during the post-EQS period. Shepherd (2004) examined the impact of the end of the EQS on price transmission from producer to international prices; and from international to retail prices employing a vector autoregression (VAR) model. Study results suggested that the EQS termination did not lead to improved price transmission because of market power exerted by coffee processors. Moreover, asymmetries in price transmission at all levels of the supply chain were identified, particularly during the post-EQS period. Gemech and Struthers (2007) examined impacts of market reforms in Ethiopia on the volatility of coffee producer prices using the Generalized Autoregressive Conditional Heteroskedasticity (GARCH) model. They found that coffee price volatility increased after market reforms, triggered by the end of the EQS. Mehta and Chavas (2008) studied the price effects of the EQS termination and found that the short-run retail prices response was greater for increases than for decreases in international prices during the post-EQS period. In contrast, they found no evidence of price transmission asymmetries between producer and international prices. More recently, Gómez, Lee and Körner (2010) examined price transmission from international to retail coffee prices by employing an asymmetric error correction model (AECM) and found evidence of short-run asymmetries with substantial differences among importing countries. This study extends and refines this literature by accounting for long-run threshold effects and short-run asymmetries in error correction models. This approach adds precision to impact evaluations for private and public decision-makers who are concerned with policies that influence the international trade of agricultural commodities.

This paper is organized as follows. We first review the literature that deals with the threshold approach to analyses of price transmission in agricultural commodity markets. Next, we develop an asymmetric threshold error correction model (ATECM) to examine impacts of the coffee EQS termination. In turn, we describe our data and present empirical results. Finally we summarize our findings and discuss the benefits and limitations of the ATECM representation to assess policies affecting price transmission in global commodity supply chains.

2.2 Modeling threshold co-integration in price transmission

A number of studies have utilized the threshold approach to examine price transmission in supply chains of agricultural commodities. Goodwin and Holt (1999) employed a threshold error correction model (TECM) to evaluate linkages between producer, wholesale, and retail prices in U.S. beef markets. Subsequently, Goodwin and Piggott (2001) developed a TECM to examine market integration among four corn and soybean markets in North Carolina, while accounting for transaction costs. More recently, Abdulai (2002) employed the threshold co-integration model developed by Enders and Granger (1998) to analyze price transmission between producer and retail prices in Switzerland's pork supply chain. He compared a standard ECM with a TECM using the Akaike and Schwarz information criteria and showed that the threshold representation is superior to its standard counterpart. Meyer (2004) considered the possibility of transaction costs in separated markets during the price transmission process with a vector error correction model following the procedures of Balke and Fomby (1997). He examined market integration between swine markets in Germany and the Netherlands and found the existence of significant transaction costs. Overall, these studies generally confirm the existence and relevance of thresholds between spatially separated markets and indicate that TECM representations generally show a faster adjustment towards the long-run equilibrium than their standard counterparts.

In this study, we follow and extend the threshold co-integration approach developed by Enders and Granger (1998) to incorporate two relevant properties in price transmission: the existence of thresholds in the co-integrating vector and the possible asymmetries in short-run price responses. As Balke and Fomby (1997) pointed out, the conventional co-integration tests of Johansen (1992a, 1992b, 1995) may be misspecified if there are thresholds in the adjustment to the long-run equilibrium. To overcome this problem, Enders and Granger (1998) suggested an alternative to the standard augmented Dickey-Fuller (ADF) regression model.

Consider two price series, RP_t retail coffee prices and IP_t international coffee prices at time period *t*. Both price variables are assumed to be integrated of order one, *I(1)*. If these two price series are co-integrated, the error term $\varepsilon_t = RP_t - \sigma_0 - \sigma_1 IP_t$ indicates deviations from the long-run equilibrium. In the presence of threshold effects, Enders and Granger (1998) proposed the following threshold autoregressive (TAR) representation of ε_t:

$$\Delta\varepsilon_t = I_t\left[\rho_0^{OUT} + \rho_1^{OUT}\varepsilon_{t-1}\right] + (1-I_t)\left[\rho_0^{IN} + \rho_1^{IN}\varepsilon_{t-1}\right] + \sum\nolimits_{i=1}^{p-1}\gamma_i\Delta\varepsilon_{t-1} + v_t, \tag{2.1}$$

where the autoregressive process of ε_t can be separated into two regimes, namely the *'OUT'* and the *'IN'* regime, depending on whether or not the threshold variable $\left|\varepsilon_{t-d}\right|$ exceeds a threshold value θ. The *'OUT'* regime represents deviations outside the threshold interval $\left[-\theta, \theta\right]$; and the *'IN'* regime represents deviations of magnitude smaller than the threshold θ (i.e. the inside threshold interval). In equation (2.1), the Heaviside indicator function, I_t, is defined as $I_t = 1$ if $\left|\varepsilon_{t-d}\right| > \theta$, zero otherwise. The parameter d represents the delay in the change from one regime to the other (e.g. $d=1$ indicates a change on an immediate occurrence) and is typically determined through statistical testing (Goodwin and Holt 1999; Gouveia and Rodrigues 2004).

Tsay (1989) suggested a nonparametric approach to identify the existence of threshold effects in the autoregressive process of variable (i.e. error correction term ε_t in this case). Following Tasy's approach, we employ a recursive least squares procedure to examine whether or not the coefficients of this autoregressive process are constant. Subsequently, we compute *TAR-F* statistics to test the null hypothesis of no changes in the parameter estimates of the autoregressive representation of ε_t. Rejection of the null hypothesis suggests the existence of a threshold θ, which must be estimated using the procedure explained below (see the details in Tsay 1989).

If threshold effects in the autoregressive series of ε_t are observed, we estimate the unknown threshold value θ using Chan's (1993) grid search method in which threshold values are obtained through a search over all possible threshold values minimizing the Squared Sum of Errors (SSE). Specifically, the threshold variable $|\varepsilon_{t-d}|$ is first sorted from the lowest to the highest value. Second, the *TAR* model in equation (2.1) is estimated using the ordered values of $|\varepsilon_{t-d}|$ as thresholds. Finally the threshold value θ is chosen so that the SSE is minimized. Hansen (1997) argued that the conventional test is not appropriate because the null hypothesis of linearity in the *AR* process does not follow a standard distribution. Consequently, he proposed a Chow-type test for threshold values using simulation methods, and provided asymptotic *p*-values based on bootstrap methods (Hansen 1997; Goodwin and Holt 1999; Goodwin and Piggott 2001).

Once the presence of threshold effects in the autoregressive process of ε_t is confirmed, the threshold error correction model (TECM) can be estimated. International and retail prices may be determined simultaneously given that these three countries account for about 50 percents of total imports from all origins (ICO 2010). Therefore, we employ Seemingly Unrelated Regression (SUR) estimation (Zellner 1962) taking into account threshold effects in the error correction process in long-run and short-run dynamics in price transmission. We also incorporate identification variables for short-run dynamics in the retail and international price equations. These variables allow us to control for additional factors that may influence retail prices (e.g. exchange rates) and international prices (e.g. precipitation in coffee producing region). A simultaneous representation of the system of equations yields:

$$\Delta RP_t = \alpha_0 + \alpha_1^{OUT} I_t \hat{\varepsilon}_{t-1}^{OUT} + \alpha_1^{IN}(1-I_t)\hat{\varepsilon}_{t-1}^{IN} + \sum_{i=1}^{p} \alpha_{2,i} \Delta RP_{t-i} + \sum_{i=0}^{p} \alpha_{3,i} \Delta IP_{t-i} + \sum_{i=0}^{p} \alpha_{4,i} \Delta z_{1,t-i} + u_{1,t} \quad (2.2)$$

$$\Delta IP_t = \beta_0 + \beta_1^{OUT} I_t \hat{\varepsilon}_{t-1}^{OUT} + \beta_1^{IN}(1-I_t)\hat{\varepsilon}_{t-1}^{IN} + \sum_{i=1}^{p} \beta_{2,i} \Delta IP_{t-i} + \sum_{i=0}^{p} \beta_{3,i} \Delta RP_{t-i} + \sum_{i=0}^{p} \beta_{4,i} \Delta z_{2,t-i} + u_{2,i} \quad (2.3)$$

where $\Delta z_{k,t-i}\ \forall k = 1,2$ are the identification variables. The term $\hat{\varepsilon}_{t-1}^{OUT}$ represents deviations from the long-run equilibrium larger than the absolute value of the threshold θ. The term $\hat{\varepsilon}_{t-1}^{IN}$ represents deviations from the long-run equilibrium within the threshold interval $[-\theta, \theta]$.

Possible short-run asymmetries in price transmission can be examined by splitting the variables ΔRP_{t-i}, ΔIP_{t-i} and $\Delta z_{1(2),t-i}$ into positive and negative changes (von Cramon-Taubadel and Loy 1996). As a result, an asymmetric threshold error correction model (ATECM) can be specified by modifying equations (2.2) and (2.3) as follows:

$$\begin{aligned}\Delta RP_t = {} & \alpha_0 + \alpha_1^{OUT} I_t \hat{\varepsilon}_{t-1}^{OUT} + \alpha_1^{IN}(1-I_t)\hat{\varepsilon}_{t-1}^{IN} + \sum_{i=1}^{p} \alpha_{2,i}^{+} \Delta^{+} RP_{t-i} + \sum_{i=1}^{p} \alpha_{2,i}^{-} \Delta^{-} RP_{t-i} + \sum_{i=0}^{p} \alpha_{3,i}^{+} \Delta^{+} IP_{t-i} \\ & + \sum_{i=0}^{p} \alpha_{3,i}^{-} \Delta^{-} IP_{t-i} + \sum_{i=0}^{p} \alpha_{4,i}^{+} \Delta^{+} z_{1,t-i} + \sum_{i=0}^{p} \alpha_{4,i}^{-} \Delta^{-} z_{1,t-i} + u_{1,t}\end{aligned} \quad (2.4)$$

$$\begin{aligned}\Delta IP_t = {} & \beta_0 + \beta_1^{OUT} I_t \hat{\varepsilon}_{t-1}^{OUT} + \beta_1^{IN}(1-I_t)\hat{\varepsilon}_{t-1}^{IN} + \sum_{i=1}^{p} \beta_{2,i}^{+} \Delta^{+} IP_{t-i} + \sum_{i=1}^{p} \beta_{2,i}^{-} \Delta^{-} IP_{t-i} + \sum_{i=0}^{p} \beta_{3,i}^{+} \Delta^{+} RP_{t-i} \\ & + \sum_{i=0}^{p} \beta_{3,i}^{-} \Delta^{-} RP_{t-i} + \sum_{i=0}^{p} \beta_{4,i}^{+} \Delta^{+} z_{2,t-i} + \sum_{i=0}^{p} \beta_{4,i}^{-} \Delta^{-} z_{2,t-i} + u_{2,t}\end{aligned} \quad (2.5)$$

where the variables $\Delta^{+} y_{t-i} = \Delta y_{t-i}$ if $\Delta y_{t-i} \geq 0$, zero otherwise; and $\Delta^{-} y_{t-i} = \Delta y_{t-i}$ if $\Delta y_{t-i} < 0$, zero otherwise, where $y = (RP, IP, z_1, and\ z_2)$.

In the threshold error correction model, price adjustments may occur only when the benefits offset the cost of changing prices due to the presence of transaction costs or other sources of price frictions (Balke and Fomby 1997). That is, the error correction mechanism operates only when deviations from the long-run equilibrium exceed a critical range [-θ, θ]. Here, the inside regime between -θ and θ can be defined as a "neutral band" within which no adjustments take place given an exogenous shock (Goodwin & Piggott 2001; Meyer 2004; Meyer & von Cramon-Taubadel 2004).

We follow a systematic approach to determine the appropriate model specification for understanding changes in price transmission in the post-EQS period. We first investigate the time series properties of international and retail coffee prices including nonstationarity and co-integration using various unit-root tests and the Johansen's co-integration test. Second, we examine possible threshold effects in the autoregressive series of the error correction term, ε_t in equation (2.1) following Tsay (1989). If threshold effects are found, we then estimate the threshold value θ using Chan's (1993) grid search method; and we test for the statistical significance of the threshold estimates following Hansen (1997). Third, for each importing country, we estimate the system of equations (2.2)-(2.3) for a symmetric TECM and the system of equations (2.4)-(2.5) for an asymmetric TECM for two periods: The EQS period, from January 1980 through December 1989; and the post-EQS period, from January 1990 to December 2009. Next, we test the short-run parameter asymmetries using *F*-tests under the null hypothesis of symmetries and employ the AIC model selection criteria to assess whether a symmetric or an asymmetric representation is more appropriate to compare price transmission behavior between the EQS and the post-EQS period.

2.3 Data

We employ monthly data of international coffee prices (the weighted average price of different coffee bean varieties) and retail prices of roasted coffee in France, Germany and the United States during the period January 1980 to December 2009. These data were obtained from the International Coffee Organization (ICO). Retail prices of roasted coffee and international prices are denoted in US dollars per pound. For identification purposes, we compile monthly exchange rates of the French Franc and the German Mark[15] to the US dollar for France and Germany and the Consumer Price Index for food and beverage for the United States. These data are from the Federal Reserve Bank Statistics (2010) and the Bureau of Labor Statistics (2010). Moreover, monthly average precipitation observed in Fortaleza, Brazil is used for the international price equation since weather patterns in this region influence international coffee prices. Precipitation data were compiled from the National Centre for Atmospheric Research (2010).

In Table 2.1, we present descriptive statistics of these data. The results show that the average retail price in Germany ($4.13 per pound) is the highest among the three countries. The coefficients of variation in Table 2.1 suggest that international prices are more volatile than retail prices in the three countries. This suggests that changes in international prices may not be fully and instantaneously transmitted to retail prices. Retail prices in the United States exhibit less variation (coefficient of variation equals 0.16) than retail prices in France and Germany (coefficients of variation equal 0.22 and 0.20, respectively).

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[15] Conversion factors for the Franc and the Mark and Euro were employed since January 2002. For German Marks, 1 Euro = 1.95583 DM; for French Francs, 1 Euro = 6.55957FF. Here we assume full transmission of the exchange rates to retail prices, since retail prices in France and Germany are denominated in dollars.

Table 2.1 Descriptive statistics of data, 1980:1-2009:12 (N=360)

	Mean	Std. Dev.	C.V.[b]	Max	Min
International price	1.014	0.365	0.360	2.042	0.412
Retail price in France	3.061	0.674	0.220	4.717	1.904
Retail price in Germany	4.125	0.810	0.196	6.179	2.473
Retail price in the US	3.136	0.510	0.163	4.669	2.352
Exchange Rate (Franc/US Dollar)	5.982	1.171	0.196	4.041	10.093
Exchange Rate (Mark/US Dollar)	1.861	0.420	0.226	3.303	1.241
Consumer Price Index, Foods and Beverages [a]	1.480	0.373	0.252	0.833	2.192
Precipitation (100mm)	1.348	1.527	1.133	8.310	0

a. Index 2000 = 1.

b. Coefficient of Variation.

2.4 Results

2.4.1 Test of integration and co-integration

We first test the time-series properties of the price data during the EQS and the post-EQS periods. We conduct the augmented Dickey-Fuller (*ADF*) and *DF-GLS* (Elliott, Rothenberg and Stock 1996; Elliott 1999) tests for the null hypothesis of nonstationarity; and we also use the *KPSS* (Kwiatkowski et al. 1992) test for the null hypothesis of stationarity. We present test results for the levels and the first differences of each price series in Table 2.2. The *ADF-t* and *DF-GLS* tests for all first difference variables (international price and retail prices in the three countries) indicate rejection of the null hypothesis of nonstationarity. Furthermore, the *KPSS* test do not reject the null hypothesis of stationarity, showing that all price series in first differences follow *I(0)* processes.

Table 2.2 Integration tests of price series

Variables in Levels			Retail Price in France	Retail Price in Germany	Retail Price in the U.S.	Inter. Price
ADF-t	$H_0 \sim I(1)$	EQS	-1.87	-2.50	-2.76	-2.40
		Post- EQS	-1.49	-1.68	-2.69	-2.47
DF-GLS	$H_0 \sim I(1)$	EQS	-0.92	-1.27	-0.47	-1.29
		Post- EQS	-0.66	-0.01	-0.13	-0.18
KPSS	$H_0 \sim I(0)$	EQS	0.63** [a]	0.34*	0.61**	0.45*
		Post- EQS	0.69**	0.80**	0.56**	0.41*
Variables in First Differences						
ADF-t	$H_0 \sim I(1)$	EQS	-7.13**	-7.96**	-7.30**	-7.64**
		Post- EQS	-12.47**	-11.89**	-10.32**	-13.21**
DF-GLS	$H_0 \sim I(1)$	EQS	-7.15**	-7.95**	-7.32**	-7.63**
		Post- EQS	-9.61**	-10.87**	-10.32**	-13.21**
KPSS	$H_0 \sim I(0)$	EQS	0.31	0.26	0.10	0.08
		Post- EQS	0.17	0.14	0.04	0.06

a. ** and * indicate 5% and 10% significant level, respectively.

We follow Johansen's (1992a, 1992b, 1995) approach to test whether international and retail price series are co-integrated. This procedure identifies the number of equations that determine the co-integration relationship between these prices in each country. For both periods and the three countries, we conduct λ_{max} and *trace* tests between international and retail prices. We present test results in Table 2.3, where r represents the co-integration rank (i.e. the number of co-integration vectors). Test results indicate that the relationship between international and retail prices in each country has at least one co-integrating vector. These results imply the existence of a long-run relationship between two price series in the three countries.

Table 2.3 Test of co-integration (*Johansen* test)

France	$H_0{:}r$	EQS period	Post- EQS period
λmax	0	10.43*[a]	29.56**
trace	0	11.92*	29.63**
Germany	$H_0{:}r$	EQS period	Post- EQS period
λmax	0	16.31**	16.99**
trace	0	17.74**	17.03**
United States	$H_0{:}r$	EQS period	Post- EQS period
λmax	0	26.56**	38.60**
trace	0	28.34**	38.60**

[a] ** and * indicate 5% and 10% significant level, respectively.

2.4.2 TAR parameter estimates

Table 2.4 presents the parameter estimates of the TAR model in equation (2.1). We employ the AIC and SBC criteria to identify the optimal lag structure of each TAR model. The delay parameter, *d* is selected as the lag value showing the largest *TAR-F* statistics in the *Tsay* test (Goodwin and Holt 1999; Goodwin and Piggott 2001). Interpreting the test statistics, we find evidence of threshold effects in the co-integrating vector (ε_{t-1}) in both periods (EQS and post-EQS) and in the three countries. The test statistics suggest that the null hypothesis of a linear autoregressive (*AR*) process in the co-integrating vector is rejected at the 5 percent significance level in the three countries and in both time periods. In addition, the selected delay parameters *d* for both periods show that the threshold lags, which represent the reaction time determining the switch between regimes (*'IN'* and *'OUT'* in equation (2.1)) in the error correction process is smaller in the post-EQS period - two, one and three months in France, Germany and the United States, respectively; versus six in all countries in the EQS period. This indicates that the time lag associated with regime switching becomes more instantaneous in the post-EQS period. The threshold value θ presented in Table 2.4 is estimated using Chan's grid search method. The magnitude of the thresholds decreases during the post-EQS period relative to its EQS counterpart: from 0.28 to 0.20 in France; 0.48 to 0.18 in Germany; and 0.18 to 0.08 in the United States. Following Balke and Fomby (1997), the interval $[-\theta, \theta]$ can be interpreted as the range where no price adjustments toward the long-run equilibrium occur. This may be due to transaction costs arising from adjusting retail prices in response to changes in international prices. These values suggest that the magnitude of deviations beyond which price adjustments take place decreases in the post-EQS period. That is, even smaller changes in international prices result in changes in retail prices in all three countries.

Table 2.4 TAR estimates

		France	Germany	United States
Optimal Lags (*p*)[a]	EQS	1	2	2
	Post- EQS	5	5	7
Delay Parameters (*d*)[b]	EQS	6	6	6
	Post- EQS	2	1	3
Tsay (1997) Test[c]	EQS	4.42** (0.01)	3.91** (0.01)	2.96** (0.04)
	Post- EQS	2.56** (0.02)	3.70** (0.00)	2.37** (0.02)
Hansen (1997) Test[d]	EQS	7.77** (0.00)	5.74** (0.00)	9.33** (0.00)
	Post- EQS	4.83** (0.00)	3.63** (0.03)	6.80** (0.00)
Threshold (θ)	EQS	0.28 (20.2%)[e]	0.48 (55.3%)	0.18 (31.9%)
	Post- EQS	0.20 (36.1%)	0.18 (23.8%)	0.08 (24.5%)
Long-run Asymmetry across Regimes[f]	EQS	2.72*	10.26**	23.64**
	Post- EQS	17.15**	14.80**	15.81**

a. Optimal lags are determined by *AIC* and *SBC*.

b. Delay parameters are chosen by the lags giving the largest *TAR-F* statistics from *Tsay* test.

c. The null hypothesis of *Tsay* test is that AR follows a linear process in a recursive least square estimation. That is, $\beta_{1,m-1} = 0$ in equation (1.4). The *F* test for no linear process and parenthesis shows asymptotic *p* values for test statistics.

d. The null hypothesis of *Hansen* test is that there are no threshold effects in autoregressive representation of variable. The *F* test for no threshold effects and parenthesis indicates asymptotic *p* values of bootstrap simulations with 100 replications.

e. Numbers in parenthesis indicate the percentage of observed deviations that lie inside the threshold.

f. The null hypothesis is that $\rho_1^{IN} = \rho_1^{OUT}$ in equation (1.1).

g. ** and * indicate 5 percent and 10 percent level of significance, respectively.

These results are consistent with the argument that termination of the EQS should result in more volatile prices. Retail prices in importing countries seem to be more sensitive to small changes in international prices. This reflects the fact that exporter countries lost the ability to set international prices during the post-EQS period, whereas downstream segments of supply chains in importing countries increased their influence to coordinate international and retail prices.

Table 2.4 also shows the percentage of observed deviations that lie in the inside regime (i.e. deviations from the long-run equilibrium within the interval $[-\theta, \theta]$). The results indicate that while the percentage decreases during the post-EQS period in Germany and in the United States, it increases in France. A percentage decrease implies that price adjustments take place more frequently in the post-EQS period relative to the EQS period. The German results yield the steepest decline (from 55 to 24 percent) of observations in the inside regime, suggesting a more extensive adjustment toward the long-run equilibrium after the EQS termination. In the United States the percentage of observations in the inside regime decreases modestly, from 32 percent in the EQS period to 25 percent in the post-EQS period. These results support the argument that downstream market segments wielded increasing influence over international-to-retail price spreads. On the contrary, this percentage in France increases in the post-EQS period. This result indicates that price adjustments take place less frequently in the post-EQS period than the EQS period, although retail prices become more responsive after the end of the EQS (e.g. the decreased magnitude of threshold value). Thus, in France there might be other factors generating price friction in the post-EQS period.[16]

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[16] For example, the Galland Law passed in 1996 limited the ability of processors and retailers to modify prices (Gomez, Lee and Körner 2010).

The Hansen tests presented in Table 4 also reject the null hypothesis of no threshold effects for both periods and all three countries at the 5 percent significance level. Additionally, the *F* statistic testing the null hypothesis that the autoregressive coefficients of each regime in equation (2.1) are equal ($\rho_1^{IN} = \rho_1^{OUT}$) also confirms that the coefficients are different across regimes. These tests provide additional evidence of threshold effects in the co-integrating vector of each country.

2.4.3 Model selection

Given the existence of threshold effects in the co-integrating vector of each country, we examine possible short-run asymmetries in contemporary and lagged explanatory variables for both periods. For this purpose, we first estimate an asymmetric TECM in equations (2.4)-(2.5) using SUR for the EQS and the post-EQS periods.[17] Next we test the null hypothesis that the coefficients for positive and negative changes in the independent variables in first differences are equal (e.g. $\alpha_{3,i}^{+} = \alpha_{3,i}^{-}$). In table 2.5, we show the χ^2 statistics corresponding to the null hypothesis of symmetry between positive and negative variables in contemporary and lagged international price changes for both periods (EQS and post-EQS). Our results provide evidence of symmetries during the EQS period and asymmetries during the post-EQS period in France. The null hypothesis of symmetries between split variables in contemporary and lagged international prices is accepted in the EQS period, but rejected in the post-EQS period. In contrast, there is modest evidence of asymmetries for Germany and the United States in both periods. The null

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[17] The exogeneity test results indicate that international prices are weakly exogenous to retail prices. The parameter estimates of the OLS and SUR specifications are practically the same. However, we prefer to use the SUR given the best fit according to Akaike Information Criteria.

hypotheses of symmetries are accepted for contemporary variables, but rejected for lagged variables in both periods.

Table 2.5 Tests of short-run asymmetries (Retail price equation)

Period	$\chi^2(1)$ a	Null Hypothesis	France	Germany	United States
EQS period	3.84	$\Delta^+ IP_t = \Delta^- IP_t$	0.120	0.018	0.527
		$\Delta^+ IP_{t-1} = \Delta^- IP_{t-1}$	1.576	6.117**	8.169***
Post- EQS period	3.84	$\Delta^+ IP_t = \Delta^- IP_t$	4.670**	0.736	1.603
		$\Delta^+ IP_{t-1} = \Delta^- IP_{t-1}$	15.082***	4.234**	42.484***

a. Critical value at 5% significance level.

Because the above tests tend favor asymmetric specifications (Meyer and von Cramon-Taubadel 2004), we further employ additional model selection criteria for robustness.[18] The AIC measures presented in Table 2.6 provide additional information for model selection. During the EQS period, the AIC values of the symmetric model specifications are smaller than their asymmetric model counterparts in all three countries. This indicates that a symmetric formulation (equation (2.2)-(2.3)) is more appropriate for the EQS period. In contrast, goodness-of-fit measures for the post-EQS period favor an asymmetric representation. Overall, Table 2.6

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[18] Meyer and von Cramon-Taubadel (2004) show that the standard tests of short-run asymmetries often lead excessive rejection of the null hypothesis of symmetry.

supports a symmetric TECM for the EQS period and an asymmetric TECM for the post-EQS period in the three countries.[19]

Table 2.6 Comparison of the AIC measures across model specifications[a]

	EQS period		Post- EQS period	
	Symmetric TECM	*Asymmetric TECM*	*Symmetric TECM*	*Asymmetric TECM*
France	**-1290.02**[b]	-1275.73	-2635.71	**-2667.10**
Germany	**-1168.79**	-1163.93	-2333.57	**-2337.27**
United States	**-1269.25**	-1252.46	-2327.11	**-2391.95**

a. AIC is the Akaike Information Criteria; TECM is Threshold Error Correction Model; and ATECM is Asymmetric Threshold Error Correction Model.

b. Number in bold indicates smaller value of AIC measure.

2.4.4 Price transmission in the long-run

Tables 2.7, 2.8 and 2.9 show the parameter estimates of the retail price equations (2.2) and (2.4) corresponding to a symmetric TECM model during the EQS period and an asymmetric TECM model in the post-EQS period for France, Germany and the United States, respectively.[20]

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[19] In model formulation, a linear trend is considered in the United States model for the post-EQS period. Other than France and Germany, co-integration test with a linear trend also suggests that the U.S. model in the post-EQS period have a possibility of a linear trend in co-integrating vector.

[20] We do not present the parameter estimates of the international price equation. The estimates results are available from authors upon request.

Table 2.7 Estimation results for France (Retail price equation)

Variables	EQS period (*Symmetric TECM*)		Post- EQS period (*Asymmetric TECM*)	
	Parameter Estimates	Standard Errors	Parameter Estimates	Standard Errors
Constant	- 0.001	(0.004)	- 0.001	(0.007)
ε_{t-1}^{OUT}	- 0.048***[a]	(0.010)	- 0.043***	(0.009)
ε_{t-1}^{IN}	- 0.026	(0.021)	0.009	(0.026)
ΔRP_{t-1}	0.500***	(0.074)	--	
$\Delta^{+} RP_{t-1}$	--		0.576***	(0.063)
$\Delta^{-} RP_{t-1}$	--		0.036	(0.106)
ΔIP_{t}	- 0.130	(0.049)	--	
$\Delta^{+} IP_{t}$	--		- 0.008	(0.053)
$\Delta^{-} IP_{t}$	--		0.239***	(0.084)
ΔIP_{t-1}	- 0.003	(0.049)	--	
$\Delta^{+} IP_{t-1}$	--		0.230***	(0.057)
$\Delta^{-} IP_{t-1}$	--		- 0.196**	(0.078)
$\Delta Exrate_{t}$	- 0.461***	(0.020)	--	
$\Delta^{+} Exrate_{t}$	--		- 0.559***	(0.044)
$\Delta^{-} Exrate_{t}$	--		- 0.423***	(0.039)
$\Delta Exrate_{t-1}$	0.179***	(0.042)	--	
$\Delta^{+} Exrate_{t-1}$	--		0.013	(0.067)
$\Delta^{-} Exrate_{t-1}$	--		0.246***	(0.046)
R^2	0.85		0.77	
N	120		240	

a. *** and ** indicate 1% and 5% significant level, respectively.

Table 2.8 Estimation results for Germany (Retail price equation)

Variables	EQS period (*Symmetric TECM*)		Post- EQS period (*Asymmetric TECM*)	
	Parameter Estimates	Standard Errors	Parameter Estimates	Standard Errors
Constant	- 0.010	(0.007)	0.006	(0.015)
ε_{t-1}^{OUT}	- 0.062***[a]	(0.016)	- 0.046***	(0.012)
ε_{t-1}^{IN}	- 0.038	(0.024)	- 0.045	(0.075)
ΔRP_{t-1}	0.135	(0.120)	--	
$\Delta^{+} RP_{t-1}$	--		0.250**	(0.101)
$\Delta^{-} RP_{t-1}$	--		0.017	(0.094)
ΔIP_{t}	0.069	(0.081)	--	
$\Delta^{+} IP_{t}$	--		0.372***	(0.110)
$\Delta^{-} IP_{t}$	--		0.571***	(0.172)
ΔIP_{t-1}	0.127	(0.084)	--	
$\Delta^{+} IP_{t-1}$	--		0.127	(0.125)
$\Delta^{-} IP_{t-1}$	--		- 0.358**	(0.167)
$\Delta Exrate_{t}$	- 1.415***	(0.100)	--	
$\Delta^{+} Exrate_{t}$	--		- 2.464***	(0.304)
$\Delta^{-} Exrate_{t}$	--		- 1.405***	(0.278)
$\Delta Exrate_{t-1}$	0.140***	(0.202)	--	
$\Delta^{+} Exrate_{t-1}$	--		- 0.463	(0.372)
$\Delta^{-} Exrate_{t-1}$	--		0.113***	(0.309)
R^2	0.69		0.55	
N	120		240	

a. *** and ** indicate 1% and 5% significant level, respectively.

Table 2.9 Estimation results for the Unites States (Retail price equation)

Variables	EQS period (*Symmetric TECM*)		Post- EQS period (*Asymmetric TECM*)	
	Parameter Estimates	Standard Errors	Parameter Estimates	Standard Errors
Constant	- 0.002	(0.012)	- 0.082	(0.016)
trend	--		0.0002**	(0.000)
ε_{t-1}^{OUT}	- 0.044	(0.016)	- 0.119***[a]	(0.024)
ε_{t-1}^{IN}	- 0.243***	(0.052)	- 0.150**	(0.053)
ΔRP_{t-1}	0.522***	(0.078)	--	
$\Delta^{+} RP_{t-1}$	--		0.184***	(0.059)
$\Delta^{-} RP_{t-1}$	--		-0.031	(0.135)
ΔIP_{t}	- 0.575***	(0.080)	--	
$\Delta^{+} IP_{t}$	--		0.243***	(0.095)
$\Delta^{-} IP_{t}$	--		-0.049	(0.147)
ΔIP_{t-1}	0.447***	(0.088)	--	
$\Delta^{+} IP_{t-1}$	--		0.902***	(0.111)
$\Delta^{-} IP_{t-1}$	--		- 0.503***	(0.152)
$\Delta CPIFB_{t}$	0.820	(2.365)	--	
$\Delta^{+} CPIFB_{t}$	--		1.151	(1.704)
$\Delta^{-} CPIFB_{t}$	--		0.371	(6.828)
$\Delta CPIFB_{t-1}$	0.129	(2.321)	--	
$\Delta^{+} CPIFB_{t-1}$	--		0.371	(1.688)
$\Delta^{-} CPIFB_{t-1}$	--		1.960	(6.724)
R^2	0.41		0.55	
N	120		240	

a. *** and ** indicate 1% and 5% significant level, respectively.

The estimated coefficients of $\hat{\varepsilon}_{t-1}^{OUT}$ and $\hat{\varepsilon}_{t-1}^{IN}$ describe the speed of adjustment towards the long-run equilibrium in each regime where the '*OUT*' regime refers to deviations outside the threshold values [$-\theta$, θ] and the '*IN*' regime represents deviations of magnitude smaller than the threshold.

For France and Germany, the estimated coefficients of the outside regime are negative, as predicted by theory, and statistically significant in both periods. Our results indicate that, in both countries, the speed of adjustment in the outside regime decreased slightly during the post-EQS period. In France (Germany), deviations from the long-run equilibrium adjust at a rate of 0.048 (0.062) in the EQS period; whereas this speed decreases to a rate of 0.043 (0.046) during the post-EQS period. Additionally, the parameter estimates suggest that the long-run parameters for the inside regime ($\hat{\varepsilon}_{t-1}^{IN}$) in both countries are not significant. This is consistent with the existence of friction in the price transmission process for deviations below the threshold value.

In the United States, the speed of adjustment of the outside regime is negative and significant for the post-EQS period, while this parameter estimate for the EQS period is not significant. In contrast to France and Germany, the estimated coefficients of error correction terms for the inside regime are significant for both periods and the speed of adjustment in this regime is reduced to a large extent in the post-EQS period (adjusting at a rate of 0.243 to 0.138). These results may be due to the smaller deviations from the long-run equilibrium in the United States compared to France and Germany. As mentioned in TAR estimates results, the threshold values of the United States for the EQS period and the post-EQS period are 0.18 and 0.08, respectively, which are much smaller than those of the two European countries. That means, in the United States, price adjustment toward the long-run equilibrium takes place for even small deviations from the equilibrium. Additionally, the parameter estimate of the trend variable in Table 2.9 is positive and significant in the post-EQS period, indicating that international-retail price spreads have increased gradually over time.

We find fundamental differences between the EQS and the post-EQS periods in terms of the price adjustment toward the long-run equilibrium. The speed of adjustment appears to decrease during the post-EQS period in all three countries. These results are consistent with the expected changes in pricing behaviors after the end of the EQS. Under the EQS, international prices were relatively stable, and more predicable to international buyers, because the export quota was set in anticipation. In contrast, international prices were more volatile and probably less predictable during the post-EQS period. Therefore, distributors in importing countries had less information during the post-EQS period, and consequently the speed of adjustment of retail prices toward the long-run equilibrium relationship decreases.

2.4.5 Price transmission in the short-run

Tables 2.7-2.9 also show estimate results of the short-run dynamics of price transmission. These results suggest differences between periods (EQS and post-EQS) and across countries. Rather than interpreting the estimated coefficients, we use them to construct Nonlinear Impulse Response Functions (NIRFs) that reflect the short-run asymmetries in the price transmission process.

Following Potter (1995), the effect of exogenous shocks on the time path of responses depend on both the magnitude and direction of the shocks, making standard linear response functions unsuitable for this analysis (Goodwin and Holt 1999; Abdulai 2002). To overcome this problem, Potter (1995) suggested a modified representation of the linear impulse response function that involved replacing the linear predictor with a conditional expectation, such that:

$$NIRF_n(\delta; X_t, X_{t-1}, \cdots) = E\left[x_{t+n} \middle| X_t = x_t + \delta, X_{t+1} = x_{t+1}, \cdots\right] - E\left[x_{t+n} \middle| X_t = x_t, X_{t+1} = x_{t+1}, \cdots\right], \quad (2.6)$$

where X_t is a vector of observed variables and δ is the postulated impulse. In Figure 1.3, we present retail price responses to positive and negative shocks in international prices with a magnitude of one standard deviation, for the three countries. The impulse response paths are calculated for the EQS and the post-EQS period for comparative purposes. Our results show interesting differences across periods and countries. During the EQS period, international price shocks exhibit symmetric effects on retail prices in all three countries regardless of the direction of the changes in international prices. This result supports our model selection (a symmetric TECM in the EQS period; and an asymmetric TECM in the post-EQS period). However, the NIRFs suggest differences in the price transmission time path between the two European countries and the United States during the EQS period. Panel A and B of Figure 2.3 show that international price shocks are not transmitted instantly to retail prices in France and Germany. The effect of the shock in these countries lasts for about seven months. This result indicates that changes in international prices generate modest short-run impacts on retail prices in these countries, consistent with parameter estimate results in Tables 2.7 and 2.8. In contrast, panel C of Figure 2.3 suggests that, in the United States, international price shocks are completely passed on to retail prices after four months. These results support the presence of short-run retail price adjustments in response to shocks in international prices, consistent with results in Table 2.9.

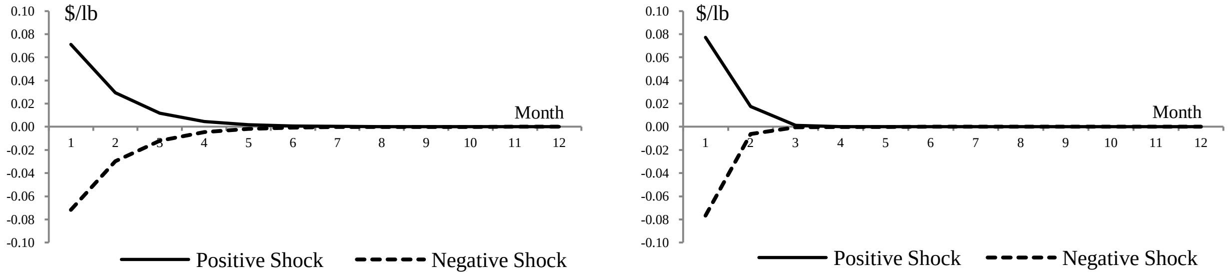
The NIRF results for the EQS period show that the U.S. domestic market is more responsive than in those of the two European countries.

Our NIRF results for the post-EQS period show how the EQS elimination impacts the short-run price transmission behaviors in these three countries. As predicted by the parameter estimate results in Tables 2.7-2.9, the NIRFs indicate that the changes in international prices have a significant short-run impact on retail prices. In particular, the responses of retail prices to international shocks seem to be asymmetric for positive and negative shocks during the post-EQS period, consistent with our model selection. However, there are differences in the time paths and their magnitude to which the shocks from the international market are completely absorbed at retail market between the two European countries and the United States. While retail prices in France and Germany tend to adjust more rapidly to negative shocks than to positive shocks, the adjustment process is quite different in the United States where positive shocks are absorbed almost entirely after two months, whereas negative shocks tend to last three months.

Our NIRFs results suggest that the end of the EQS resulted in changes in the short-run price transmission behaviors in these three countries. First, while retail prices seem to symmetrically adjust to changes in international prices in the EQS period, their responses are asymmetric in the post-EQS. Second, the effects of changes in international prices on retail prices tend to last longer in the EQS period, whereas their impacts disappear faster in the post-EQS period. These phenomena are observed in all three countries. As Bark and de Melo (1988) pointed out, the end of the EQS may have led to a rent transfer from the exporting countries to importing countries, indicating that the influence of exporting countries on international prices was reduced in the post-EQS period. This implies that, in the post-EQS period, downstream segments of the supply chain in importing countries respond more quickly and asymmetrically to changes in international prices to benefit from such changes.

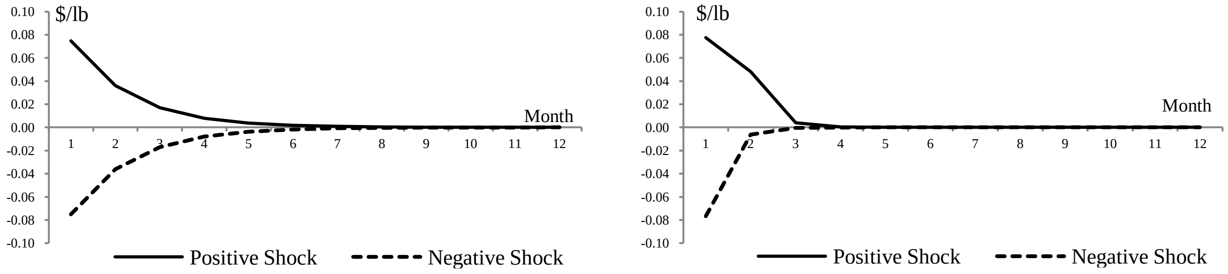
Our results also show differences between the United States and the two European countries in terms of the asymmetric price transmission behaviors in the post-EQS period. In the United States, we find evidence of Bacon’s (1991) "rockets and feathers" behavior. That is, retail prices rise faster than they fall in response to changes in international prices, suggesting that downstream segments benefit from short-run price transmission behavior. In contrast, in France and Germany, our results suggest that retail prices decrease quickly and increase slowly, contrary to the "rockets and feathers" phenomena. These differences may result from dissimilarities in market structure identified in Gómez, Lee and Körner (2010). In the United States, the coffee processing sector is highly concentrated and the level of food retail concentration is moderate. In addition, the product volume handled by private label brands and hard retail discounters is relatively small. Therefore, coffee processors may have more ability to coordinate the supply chain and benefit from short-run asymmetries. In contrast, the European coffee processing (food retailing) sector is less (more) concentrated and the share or retailer private label is higher than in the United States. Moreover, a unique characteristic of French/German coffee market is the relatively larger market share of hard-discounter retailers (e.g. Aldi). Hard discounters often employ aggressive competitive strategies based on low prices relative to competitors. Indeed, the French and German retail sectors experienced coffee price wars in the early and late 1990s, respectively (Körner 2002). Such market structure features may explain the price patterns observed in Germany and France.

A. France


<EQS Period>

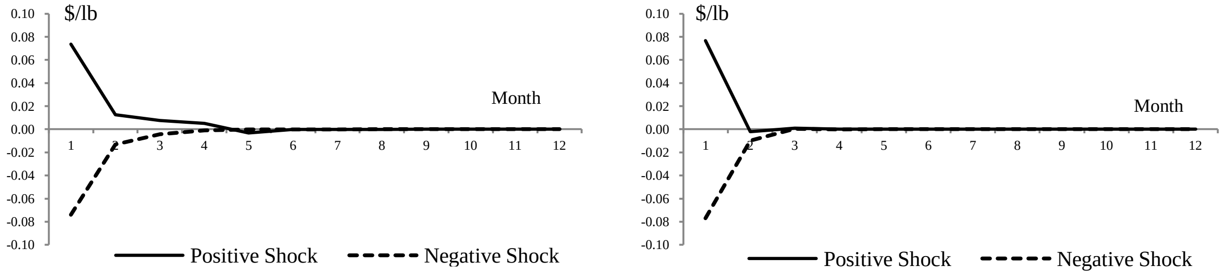
<Post-EQS period>

B. Germany


<EQS period>

<Post-EQS period>

C. United States


<EQS period>

<Post-EQS period>

Figure 2.3 Responses of retail prices to the changes in international prices

2.5 Concluding remarks

In this study we investigated price transmission between international and retail coffee prices in the three largest coffee-importing countries, France, Germany and the United States. We examined differences in price transmission behaviors between the EQS and the post-EQS periods, taking into account the existence of threshold effects and asymmetries in the price transmission process.

Our analysis uncovered fundamental differences in long-run and short-run price transmission behaviors between periods. During the post-EQS period, retail prices became more responsive to changes in international prices (i.e. the threshold values θ became smaller), because trade liberalization appears to lead to greater integration between international and retail prices. Nevertheless, the speed of adjustment appeared to decrease slightly during the post-EQS period in all three countries. A plausible explanation is that international prices became more volatile during the post-EQS period, generating more uncertainty about international prices. In these circumstances, distributors in importing countries had less information to make pricing decision during the post-EQS period. The speed of adjustment toward the long-run equilibrium slowed as a result. Our nonlinear impulse response analyses indicate that in the short-run, changes in international prices were transmitted faster and asymmetrically to retail prices during the post-EQS period in all three countries. These short-run dynamics may result from the increased influence of downstream segments of the supply chain in importing countries on pricing decision of international prices. However, our results show differences between the United States and the two European countries in terms of the nature of observed price transmission asymmetries. That is, we find evidence of "rockets and feathers" behavior in the

United States, where retail prices rise faster than they fall in response to changes in international prices. In contrast, in France and Germany, retail prices seem to fall quickly and rise slowly, which is contrary to the "rockets and feathers" phenomenon. We explained these differences in terms of dissimilarities in market structures across countries.

Our study provides valuable insights on the application of an ATECM representation for policy evaluation, but several limitations indicate the need for further research. There may be demand and supply factors other than termination of the EQS that may explain the changes in price transmission between international and retail coffee prices after 1990. In particular, price transmission from upstream to downstream markets in food supply chains is closely related to market structure. Therefore, the extent of price transmission depends on the behavior of consumers and firms as well as on the exertion of market power by supply chain participants. Consequently, future research on price transmission using threshold error correction models should incorporate formal models of market structure and their conduct. In addition, our model assumes full transmission of the exchange rates to domestic prices because domestic prices in France and Germany are denominated in dollars. Future research using error correction models to examine price transmission in global supply chains should allow for incomplete exchange rate transmission.

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CHAPTER 3

OLIGOPOLY POWER AND COST PASS-THROUGH IN THE U.S. AND THE GERMAN COFFEE MARKET

Abstract

In this study, we examine linkages between non-competitive pricing behavior and cost pass-through patterns in markets for the roasted coffee in the United States and Germany, the two largest coffee importing countries. We develop a structural supply-demand model rooted in the new empirical industrial organization (NEIO) framework. We employ an iterative generalized method of moments (GMM) for estimation to evaluate the degree of market power. Subsequently, we use a threshold error correction model (TECM) to test whether the time-dependent market power measure influences cost pass-through behavior in the two countries. We find that the roasted coffee markets in both countries are oligopolistic. The degree of market power is stronger in Germany than in the United States, supporting the market observation that the German roasted coffee supply chain is more concentrated than in the United States. We find that market power leads to an increase of the pass-through rates when international prices are rising, but it has a modest effect on pass-through rates when they are falling. These results provide evidence of "rockets and feathers" behavior in retail pricing in both countries. The "rockets and feathers" phenomenon seems to be mainly caused by the exertion of market power in the United States. In contrast, other factors different than market power seem to also lead to this conventional phenomenon in Germany.

3.1 Introduction

Non-competitive pricing behavior in markets for frequently purchased goods such as food and beverage products has been of concern to policy makers and consumers alike. In recent years, given sharp farm gate price increases for agricultural commodities, consumers often observe that food retail prices respond faster to commodity price increases than to commodity price decreases. This casts doubt on the practice of market power by food distributors (e.g. manufacturers and retailers) reflected in an asymmetric cost pass-through to retail prices. Unlike this mainstream perception, food distributors occasionally confront tradeoffs between profit and market share in their pricing decisions. That is, they may prefer setting relatively high retail prices in comparison to the equilibrium prices by exercising market power to increase their margins. However, at the same time, they may also have incentives to set retail prices relatively low to maintain or expand their market share in a competitive environment.

A substantial body of empirical literature shows that a positive or negative input price change in vertically-related food markets tends not to be completely, symmetrically and instantly passed on to downstream segments (Von Cramon-Taubadel 1998; Goodwin and Holt 1999; Abdulai 2002; Gómez, Lee and Körner 2010). As referred to in many studies, exertion of market power in an oligopolistic environment is often associated with incomplete and asymmetric cost pass-through to retail prices (Borenstein, Cameron and Gilbert 1997; Meyer and von Cramon-Taubadel 2004). Most studies use the error correction model (ECM) to study incomplete and asymmetric pass-through. However, the evidence generated by this approach is only indicative of market power because its theoretical grounding does not allow for distinguishing the source of variation in pass-through rates over time. (Digal and Ahmadi-Esfahani 2002; Meyer and von

Cramon-Taubadel 2004). In contrast, industrial organization theory offers an appropriate framework to explicitly measure the degree of market power by oligopolistic firms making pricing decision.

In this study, we shed new light on the relationship between non-competitive pricing behavior and cost pass-through patterns, focusing on the roasted coffee market in the United States and Germany, the two largest coffee importing countries.[1] During the last two decades, coffee bean prices traded in international markets have exhibited substantial volatility. Figure 2.1 illustrates movements in monthly average retail roasted coffee prices along with prices of three coffee bean varieties, *Columbian Mild*, *Other Mild* and *Robusta*, traded in New York and Germany exchange markets over the 1992 - 2009 period. *Columbian Mild* and *Other Mild* are the primary varieties consumed in these countries; *Robusta* is much cheaper than other varieties because of high crop yields and limited usage for espresso blending. Prices for *Columbian Mild* and *Other Mild* show similar evolution in both markets. However, the corresponding average retail prices for roasted coffee in each country appear to respond differently to changes in coffee bean prices. In particular, retail prices in Germany and the United States have shown different responses for upward spikes in international bean prices in the mid-1990s. For instance, while retail prices in the United States tended to respond instantly to shocks in international bean prices, those in Germany seemed react far more slowly for changes in bean prices. These differences are graphically shown for bean price shocks in 1994-1995. In addition, the degree to which changes in international bean prices are passed on to retail prices seem to be different between countries. That is, the pass-through appears to be smaller in Germany than the United States. These

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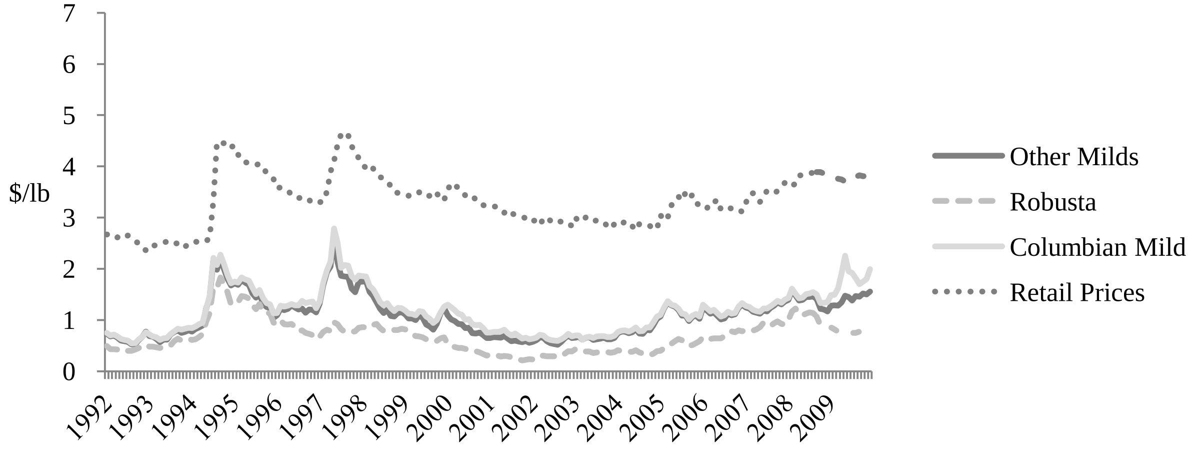
[1] United States and Germany are the largest coffee importing countries. The United States imported 1.25 million tons of green bean in 2009, which rank first among all countries. Germany ranks second, importing 1.05 million tons of green bean in same period (FAO 2011).

differences between countries may be closely related to factors associated with market structure and the exertion of market power.[2]

A. United States



B. Germany

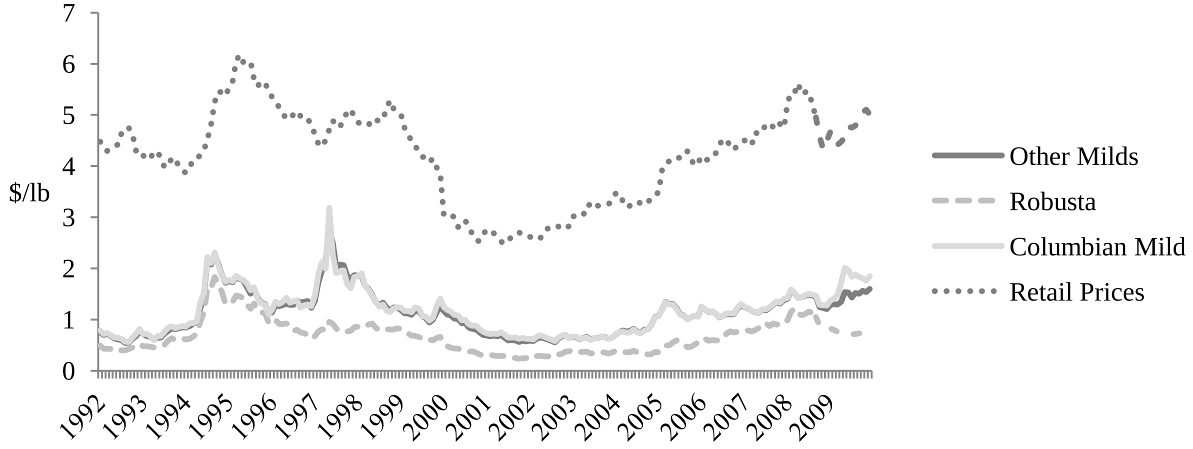


Figure 3.1 Retail roasted coffee prices and coffee bean prices, 1992:01-2009:12

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[2] Morriset (1998) pointed out that the average spread between retail coffee prices and world bean prices increased by 23 percent and 45 percent from 1975 to 1994 in Europe and the United States, respectively, which is much higher than other food commodities he studied (e.g. beef, rice, sugar and wheat). He posited that these are attributed to asymmetric responses of retail prices to upward movements and downward movements of coffee bean prices and that the market power of coffee roasting companies could be one factor in explaining increasing price spreads (Durevall 2007).

There are only a few studies on the links between market power and cost pass-through. Peltzman (2000) evaluated price transmission asymmetries on 242 consumer and producer goods and measured the degree of asymmetry. He examined correlations with proxies of imperfect competition such as the number of competitors and the degree of market concentration using the Herfindahl-Hirschman Index. He found that asymmetries were more extreme as the number of competitors falls, but they decrease when market concentration increases. Although Peltzman's (2000) approach is useful in cross-section analysis, it is not appropriate for the study of specific industries (e.g. coffee roasting) because the proxy capturing non-competitive pricing behavior does not exhibit enough variability over time (Meyer and von Cramon-Taubadel 2004). More recently, Richards, Allender and Hamilton (2011) investigated how falling and rising input prices influence market power in potato and fluid milk markets. They estimated consumer demand employing a Generalized Extreme Value Model. Subsequently, based on the derived consumer demand, they estimated a supplier pricing model using a traditional industrial organization framework. They found that for potatoes, wholesale and retail market power increases (decreases) when input prices decrease (increase). For fluid milk, findings were that market power declines when commodity prices increase but does not change with falling commodity prices, suggesting that market power of suppliers and retailers in fluid milk market weakens when price volatility is high.

Our objective is to evaluate the degree of market power exhibited in the roasted coffee market in the United States and Germany by examining critical factors determining different pricing behaviors (Iwata 1974)[1] and to examine how the exertion of market power is associated

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[1] Iwata (1974) posited that price setting power in a homogeneous product oligopoly market may first be related to differences in price elasticity associated with per capita consumption levels. Second, competitive differences may be related to different conjectural variations. Finally, differences in marginal costs lead to different price levels.

with cost pass-through patterns in these two countries. For this purpose, we build a structural supply-demand model of roasted coffee in the context of the new empirical industrial organization (NEIO) framework to develop a time-variant measure of market power in each country. Subsequently, the market conduct measures obtained from the NEIO model allow us to investigate the impact of market power on cost pass-through rates by utilizing a threshold ECM.

We contribute to the literature, by linking market power measures obtained from a structural supply-demand model grounded on industrial organization theory to a time-series based empirical cost pass-through model. In addition, from a policy perspective, we offer an empirical test for identifying sources causing incomplete and asymmetric cost pass-through in food supply chains.

The paper proceeds with a review of the U.S. and German roasted coffee markets in section 2. In section 3, we develop a theoretical model of market supply and demand in the context of the NEIO framework; and we subsequently build an empirical cost pass-through model that incorporates time-dependent market conduct measures. In section 4 we describe the data employed in estimation. In section 5 we discuss the econometric results of the structural supply-demand model and market conduct measures. Here, we also discuss the estimated cost pass-through rates results and the influence of market power in cost pass-through behavior. Section 6 concludes and suggests areas for future research.

3.2 Roasted coffee markets in the United States and Germany

A description of the coffee markets in the United States and Germany help inform the model (Table 3.1). Total coffee consumption in the United States and Germany in 2010 were 21.7 and

9.3 million bushels, respectively. The United States and Germany are the largest coffee importing countries, accounting for about 40 percent of total world imports (FAO 2011). While U.S. coffee consumption is twice as large as that of Germany, Germans consume more coffee at the per capita level. Average annual per capita coffee consumption in 2009 was 4.1kg and 6.5 kg in the United States and Germany, respectively.

Table 3.1 Selected characteristics of coffee supply chain

	U.S.[a]	Germany[b]
Consumption		
Total coffee consumption (million bushel)	21.78	9.29
Per capita consumption (kg)	4.1	6.5
Manufactures sector		
Share of four leading brands (%)	66.2	78.9
Share of private label brands (%)	10.1	13.0
Retailing sector		
Share of five leading supermarkets (%)	35.5	61.8
Share of hard-discounter retailers (%)	<2.0	34.0

a. Source: Mintel U.S. coffee report (2011)

b. Source: Mintel Germany coffee report (2004)

Industry structures vary between these two countries in terms of the degree of concentration and market characteristics. In 2003, the German roasted coffee industry was made up of 44 firms; the four biggest companies in the German roasted coffee industry and their market share were *Tchibo* with 33.3 percent, *Melitta* with 17.3 percent, *Jacobs* with 17.2 percent and *Dallmayr* with 11.1percent of Germany coffee sales. These companies had a combined market share of 78.9 percent (Mintel 2004). In 2010, the United States had 215 roasters and was slightly less concentrated than in Germany; the four biggest companies in the U.S. roasted coffee industry and their market share were *J.M. Smucker* with 32.4 percent, *Kraft Foods* with 15.7 percent, *Starbucks Coffee Company* with 12.7 percent, and *Green Mountain Coffee Roasters* with 5.4 percent. These companies had a combined market share of 66.2 percent of total U.S. coffee sales (Mintel 2011). The shares of private label brands, which are often positioned as a lower cost alternative to major brands, are 10.1 percent and 13.0 percent in the U.S. and the German markets, respectively.

Food retailing in Germany is more concentrated than in the United States. In Germany, the market share of five leading supermarkets is about 62 percent, which is about twice as large as the share of the five largest supermarkets in the United States. A striking difference between two countries is the market share of hard discounters. The market share of hard discounters (e.g. Aldi and Lidl) is quite high in Germany, at about 34 percent. In contrast, the share of hard discounters in the United States was less than 2 percent in 2010. Hard discounters offer a limited assortment of products (typically 5,000 to 6,000 stock keeping units, compared to 45,000 stock keeping units in traditional supermarkets) in large quantities, which allows them to operate extremely low-cost supply chains (Gómez, Lee and Körner 2010). Another characteristic of the German coffee market is the presence of a coffee excise tax levied on roasters at 2.19 Euros per kilogram. In contrast, there is no excise tax on coffee in the United States.

These market characteristics suggest differences in the structure of the coffee supply chains between the United States and Germany. One may speculate that these market structures may lead to differences in cost pass-through behaviors. In next section, we develop a model that describes behaviors in demand and supply sectors, and empirically evaluate the degree of market conduct. Estimated measures on market conduct are subsequently employed in the examination of links between the extent of market power and cost pass-through behaviors in both countries.

3.3 Empirical model

Our study follows two steps to examine linkages between non-competitive pricing behaviors and cost pass-through patterns in the U.S. and the German roasted coffee markets. In the first stage, we develop an empirical model to evaluate the extent to which the roasted coffee industry exercises market power in retail pricing. In the second stage, we employ the estimated time-dependent market conduct measures in a threshold error correction model of cost pass-through. Here, the variable of market conduct is incorporated into the coefficient explaining the speed of adjustment toward the long run equilibrium between international and retail coffee prices. This allows us to measure how the degree of market power influences cost pass-through behaviors in both countries.

3.3.1 Demand and supply model of roasted coffee

We start our analysis by constructing a structural supply-demand model following Bettendorf and Verboven (2000). We first specify a demand function for roasted coffee, assumed to be linear and homogenous of degree zero in prices and income. Furthermore, roasted coffee is assumed to be a homogenous good in this study.[2] Demand for roasted coffee is a function of retail price of roasted coffee, income level of households, changes in preferences over time and prices of substitute beverages. Then the aggregate demand function can be written as:

$$Q_t = \alpha_0 + \alpha_1 p_t + \alpha_2 y_t + \alpha_3 T + \alpha_4 p_t^s \,, \tag{3.1}$$

where Q_t is total consumption quantities of roasted coffee, p_t is price for roasted coffee, y_t is the income level of households, T is a linear trend capturing changes in preferences and p_t^s refers to prices of substitute goods such as tea, instant coffee and carbonated soft drinks.

Based on this consumer demand, coffee roasters are assumed to maximize profits by choosing quantity as their strategic variable. The profit of any individual firm *i* (*i=1,⋯,N*) at time *t* can be expressed as:

$$\pi_{i,t} = \frac{1}{1+\tau} p_t(Q_t) Q_{i,t} - C_{i,t}(Q_{i,t}, w_t) \,, \tag{3.2}$$

where $Q_t = \sum_i Q_{i,t}$ is total quantities sold in the market in period *t*, $Q_{i,t}$ is quantities sold by firm *i* in period *t* and τ is a value-added tax[3]. $C_{i,t}$ represents the cost function of firm *i* in period *t* where costs of firm *i* depend on its own quantity sold in period *t*, $Q_{i,t}$ and a vector of input prices

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[2] Homogeneity is assumed because of aggregate data in the empirical analysis (see e.g. Cowling and Waterson 1976; Clarke and Davies 1982). For the U.S. (Gollop and Roberts 1979; Roberts 1984) and the Netherlands (Bettendorf and Verboven 2000), homogeneity of roasted coffee is presumed as well.

[3] Due to the presence of an excise tax τ_c on roasted coffee in Germany, the term, $\tau_c \cdot Q_{i,t}$ is incorporated in the right-hand side of equation (2.2) in the German model.

in period *t*, w_t. Differentiating equation (3.2) with respect to the firm *i*'s quantity supplied, $Q_{i,t}$, the first order condition of firm *i* yields:

$$p_t(Q_t) = (1+\tau)\frac{\partial C_{i,t}(Q_{i,t}, w_t)}{\partial Q_{i,t}} + \frac{\theta_i}{\eta} p_t(Q_t)\,. \tag{3.3}$$

In equation (3.3), $\eta = -(\partial Q_t / \partial p_t)(p_t / Q_t)$ refers to the price elasticity of demand for roasted coffee and $\theta_i = (\partial Q_t / \partial Q_{i,t})(Q_{i,t} / Q_t)$ represents the conjectural variation elasticity. The conjectural variation elasticity measures firm *i*'s strategic conjecture about other firms' behavior regarding their choice of quantities supplied, and it ranges from zero to one. A value of zero indicates that there is no response of total quantities to changes in firm *i*'s output (i.e. a perfect competitive market), whereas a value of one is interpreted as a monopolistic market structure (or a perfect collusion market).

The individual supply functions can be aggregated following Appelbaum (1982) where the cost function is assumed to have the Gorman-Polar form (Bettendorf and Verboven 2000; Durevall 2007).[4] By doing so, the market supply function is obtained multiplying equation (3.3) by $Q_{i,t} / Q_t$ and aggregating over all firms. Then the market supply function yields:

$$p_t(Q_t) = (1+\tau)MC_t(w) + \frac{\theta}{\eta} p_t(Q_t)\,, \tag{3.4}$$

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[4] This function assumes that the individual cost functions are quasi-homothetic. Therefore, it implies that at the optimum, marginal costs are equal over all firms, but allowing for different cost curves across firms.

where $MC_t(w) = \sum_i (\partial C_i / \partial Q_{i,t})(Q_{i,t} / Q_t)$ represents a weighted average marginal cost of the industry and $\theta = \sum_i \theta_i (Q_{i,t} / Q_t)$ is an average industry conduct parameter. From equation (3.4), we derive the Lerner Index to measure the degree of market power. A Lerner Index equal to zero, provided evidence of a market operating in perfect competition, whereas a value of one implies that the market is monopolistic. The Lerner Index is obtained by:

$$\frac{\theta}{\eta_t} = \frac{p_t - (1+\tau)MC_t(w)}{p_t} = L_t \tag{3.5}$$

From equation (3.1), price elasticity of demand can be computed as $\eta_t = \hat{\alpha}_1 \frac{p_t}{Q_t}$, so the market supply function of equation (3.4) can be modified as:

$$p_t(Q_t) = (1+\tau)MC_t(w) + \frac{\theta}{\alpha_1} Q_t, \tag{3.6}$$

where the weighted average marginal cost of the industry, $MC_t(w)$, is inferred indirectly from the evolution of input prices (Appelbaum 1982). Given that the production of roasted coffee is relatively simple (Sutton 1991) and that input factors are used in fixed proportions, the market marginal cost can be specified as a linear function in input prices (Bettendorf and Verboven 2000) as follows:

$$MC(w) = \beta_0 + \beta_1 w_t^m + \beta_2 w_t^c + \beta_3 w_t^r + \beta_4 w_t^l + \beta_5 w_t^d\ , \tag{3.7}$$

where w_t^m, w_t^c and w_t^r refer to prices of the three different coffee bean varieties, namely *Other Mild*, *Colombian Mild* and *Robusta*, respectively; w_t^l is labor costs in the manufacturing sector and w_t^d represents distribution costs. Using equations (3.1), (3.6) and (3.7), a structural demand and supply model for the roasted coffee market yields:

$$Q_t = \alpha_0 + \alpha_1 p_t + \alpha_2 y_t + \alpha_3 T + \alpha_4 p_t^s \qquad (3.8)$$

$$p_t = (1+\tau)\left[\beta_0 + \beta_1 w_t^m + \beta_2 w_t^c + \beta_3 w_t^r + \beta_4 w_t^l + \beta_5 w_t^d\right] + \frac{\theta}{\alpha_1} Q_t . \qquad (3.9)$$

We employ the iterative generalized method of moments (GMM) (Hansen 1982; Hansen and Singleton 1982) for estimation, given that the parameters in the system of equations (3.8)-(3.9) are nonlinear. This is a convenient approach to estimating nonlinear models without complete knowledge of the probability distribution of the data (Hall 1993; Ogaki 1993).[5] Estimating equations (3.8) and (3.9) provides parameter estimates of the conjectural variation elasticity, $\hat{\theta}$ and of the time-varying price elasticity of demand, $\hat{\eta}_t = \hat{\alpha}_1 \frac{p_t}{Q_t}$. Subsequently, the Lerner Index is derived using equation (3.5) to measure the degree of market power in the two countries.

3.3.2 Cost pass-through

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[5] The two step and the iterative GMM estimators have asymptotically same distribution. However, the iterative estimator has superior properties in finite samples (Hall 2005).

The primary variations in roasted coffee processing costs result from changes in international coffee bean prices. Therefore, we evaluate cost pass-through behaviors by analyzing how fast changes in international coffee bean prices are passed on to retail prices in domestic markets.

To do this, we employ a threshold error correction model (TECM) representation, given that two price series at different segments in a food supply chain are typically co-integrated (Von Cramon-Taubadel and Loy 1996). Particularly, the ECM approach is useful in cost pass-through analysis because the pass-through rates are easily obtained through the long-run error correction term in the ECM representation. The coefficient of the error correction term can be interpreted as the speed at which retail prices adjust to the long-run equilibrium after a shock in international coffee bean prices. Moreover, the threshold approach in the ECM allows one to take into account the price rigidities that may affect price transmission behavior and cost pass-through (Tong 1983; Balke and Fomby 1997). The existence of price frictions such as transaction costs and menu costs may lead to different pass-through rates along the food supply chain, depending on the magnitude and sign of the upstream shock (Ben-Kaabia and Gil 2007). By allowing threshold effects in the error correction process, we can identify regime-specific pass-through behaviors. Specifically, we focus on whether pass-through rates are different when international coffee prices are rising and falling. For this purpose, we follow Hansen (1999) and Lo and Zivot (2001) to consider the time series nature of price data (e.g. integration and co-integration) and the regime-specific nonlinear pass-through by employing the threshold error correction model (TECM).

We start our analysis by considering the co-integration relationship between two price series: retail prices of roasted coffee, p_t, and international coffee prices, w_t.[6] If each price series is integrated of order one, *I(1)*, and the two price series are co-integrated, then the error correction term, $\hat{\varepsilon}_t$ can be calculated as the residual from the linear relationship between p_t and w_t. The error correction term is interpreted as the deviation from the long-run equilibrium between p_t and w_t. But the adjustment to restore the long-run equilibrium may depend on the magnitude and the sign of $\hat{\varepsilon}_t$. Therefore, to enable for regime-specific nonlinear pass-through, the first step is to specify a threshold autoregressive representation (TAR) of this residual, $\hat{\varepsilon}_t$ as follows (Balke and Fomby 1997; Enders and Granger 1998):

$$\Delta\hat{\varepsilon}_t = \begin{cases} \sigma_0 + \sigma_1^{(I)}\hat{\varepsilon}_{t-1} + \sum_{i=1} \phi_i \Delta\hat{\varepsilon}_{t-i} + e_t & if\ \hat{\varepsilon}_{t-d} < \gamma_1 \\ \sigma_0 + \sigma_1^{(II)}\hat{\varepsilon}_{t-1} + \sum_{i=1} \phi_i \Delta\hat{\varepsilon}_{t-i} + e_t & if\ \gamma_1 \le \hat{\varepsilon}_{t-d} < \gamma_2 \\ \sigma_0 + \sigma_1^{(III)}\hat{\varepsilon}_{t-1} + \sum_{i=1} \phi_i \Delta\hat{\varepsilon}_{t-i} + e_t & if\ \hat{\varepsilon}_{t-d} \ge \gamma_2 \end{cases}, \tag{3.10}$$

where the regime-dependent autoregressive (AR) process of the residual $\hat{\varepsilon}_t$ is determined by the unknown threshold values γ_1 and γ_2. In equation (3.10), *d* refers to a delay parameter that explains the time lag capturing threshold behavior in the autoregressive process of $\hat{\varepsilon}_t$. The coefficients $\sigma_1^{(I)}$, $\sigma_1^{(II)}$ and $\sigma_1^{(III)}$ are determined by whether the threshold variables, $\hat{\varepsilon}_{t-d}$ are

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[6] International (composite) price is a weighted average price of coffee bean varieties traded in international market. This price is calculated and published by the International Coffee Organization (ICO).

greater than the upper threshold value, γ_2, are less than the bottom threshold value, γ_1 or are between them.

Formulating a TECM requires several steps. In the first step, we examine the existence of the threshold effects in the autoregressive process of variable, $\hat{\varepsilon}_t$, following Tsay (1989). We employ a recursive least squares procedure to examine whether or not the coefficients of this autoregressive process are linear. Subsequently, we compute *TAR-F* statistics to test the null hypothesis of no changes in the parameter estimates of the autoregressive representation of $\hat{\varepsilon}_t$. Rejection of the null hypothesis suggests the existence of a threshold effect (Tsay 1989; Goodwin and Holt 1999; Goodwin and Piggott 2001).

In the second step, we estimate the threshold values using the two-dimensional grid search method and examine the significance of the estimated threshold parameters following Hansen (1999) and Lo and Zivot (2001). We first consider two regimes defined by the threshold value, γ_1. This first TAR model includes two regimes ($\hat{\varepsilon}_{t-d} > \gamma_1$ and $\hat{\varepsilon}_{t-d} \leq \gamma_1$) and is estimated using conditional ordinary least squares. Let $S_1(\gamma_1)$ be the sum of squared errors of the two-regime TAR model. The threshold value is obtained as the value that minimizes the sum of squared errors as follows (Chan 1993; Hansen 1999):

$$\hat{\gamma}_1 = \arg\min_{\gamma_1} S_1(\gamma_1)\,. \tag{3.11}$$

Hansen (1999) suggests a method for testing whether the threshold parameter estimated in equation (3.11) is statistically significant. The likelihood ratio (*LR*) test for the null hypothesis of no threshold effect is based on the statistic $F_1 = (S_0 - S_1(\hat{\gamma}_1)) / \hat{\sigma}^2$, where S_0 represents the sum of squared errors obtained from a model with no threshold. Since the threshold value, γ_1 is not identified in the null hypothesis, the classical tests of F_1 do not follow standard distributions. To overcome this limitation, Hansen(1996) suggests a bootstrap method to calculate the *p*-value to conduct the *LR* test explained above.

We also examine the possible existence of a second threshold value in the TAR model as indicated in equation (3.10). Similar to the single threshold case, the TAR model with three regimes is estimated conditional on the fixed value $\hat{\gamma}_1$, where $\hat{\gamma}_1 < \gamma_2$. Given the first threshold $\hat{\gamma}_1$, the optimal threshold value of γ_2 can be estimated as:

$$\tilde{\gamma}_2 = \arg\min_{\gamma_2} S_2(\hat{\gamma}_1, \gamma_2), \tag{3.12}$$

where $S_2(\hat{\gamma}_1, \gamma_2)$ is the conditional sum of squared errors from the three-regime TAR model. Similarly, the *LR* test statistic for the null hypothesis of single threshold versus two thresholds can be calculated as $F_2 = \{S_1(\hat{\gamma}_1) - S_2(\hat{\gamma}_1, \tilde{\gamma}_2)\} / \tilde{\sigma}^2$. To test the statistical significance of $\tilde{\gamma}_2$, we use the same bootstrapping method described above. If F_2 is sufficiently large, the null hypothesis of a single threshold model is rejected in favor of a two-threshold model specification. The same procedure can be applied to test for additional thresholds.

If the presence of threshold effects in the error correction process is confirmed, then the threshold error correction model (TECM) representation yields:[7]

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[7] The erogeneity test results indicate that international coffee bean prices, w_t is weakly exogenous to retail roasted coffee prices, p_t in both countries.

$$\Delta p_t = \begin{cases} \lambda_0 + \lambda_1^{(I)}\hat{\varepsilon}_{t-1} + \lambda_2 \Delta p_{t-1} + \lambda_3 \Delta w_t + \lambda_4 \Delta w_{t-1} + u_t, & if \ \hat{\varepsilon}_{t-d} < \gamma_1 \\ \lambda_0 + \lambda_1^{(II)}\hat{\varepsilon}_{t-1} + \lambda_2 \Delta p_{t-1} + \lambda_3 \Delta w_t + \lambda_4 \Delta w_{t-1} + u_t, & if \ \gamma_1 \le \hat{\varepsilon}_{t-d} < \gamma_2 \ , \\ \lambda_0 + \lambda_1^{(III)}\hat{\varepsilon}_{t-1} + \lambda_2 \Delta p_{t-1} + \lambda_3 \Delta w_t + \lambda_4 \Delta w_{t-1} + u_t, & if \ \hat{\varepsilon}_{t-d} \ge \gamma_2 \end{cases} \tag{3.13}$$

where the parameter estimates of the error correction terms λ_1^k (*k=I, II, III*) are estimated for three different regimes, according to the magnitude of the threshold variable $\hat{\varepsilon}_{t-d}$. Equation (3.13) allows us to investigate how the practice of market power influences cost pass-through behavior. Recall that the pass-through rate is defined as the speed at which retail prices adjust to restore the long-run equilibrium relationship with international prices. A measure of this adjustment rate is provided by parameter estimates, λ_1^k (*k=I, II, III*) of the error correction term in each regime.

We hypothesize that the pass-through rate depends on the degree of market power. That is, if distributors (e.g. coffee roasters or retailers) in the domestic market have market power, they may adjust retail prices faster for increases and slower for decreases in international prices of green coffee by exerting market power. This cost pass-through behavior is commonly referred to as the "rockets and feathers" phenomenon (Bacon 1991). For the purpose of this analysis, the regime-specific pass-through rate parameters, $\lambda_1^{(I)}$, $\lambda_1^{(II)}$ and $\lambda_1^{(III)}$ are assumed to be linearly dependent on the time-varying Lerner Index, $\hat{L}_t$ obtained from the structural model in equations (3.8)-(3.9). Therefore, the adjustment parameters can be expressed as:

$$\lambda_1^{(k)}(\hat{L}_t) = \varphi_1^{(k)} - \varphi_2^{(k)}\hat{L}_t \, , \ \forall k = I, II, III \tag{3.14}$$

where the pass-through rate in regime *k*, $\lambda_1^{(k)}$ is expected to have an inverse relationship with the estimated time-dependent Lerner Index, $\hat{L}_t$.

By allowing regime-specific and market power-dependent pass-through rates, we can examine whether pass-through rates differ according to a rising and falling international prices. Moreover, we can test whether our measure of market power is associated with these different pass-through rates. Given the co-integration relationship $\hat{\varepsilon}_t = p_t - \rho_0 - \rho_1 w_t$ and the threshold parameters γ_1 and γ_2, the regime (*I*) ($\hat{\varepsilon}_{t-d} < \gamma_1$) is interpreted as the regime where retail prices are relatively low compared to the long-run equilibrium relationship, due to rising international prices. Therefore, in this regime, retail prices are expected to rise to reestablish the long-run equilibrium. If coffee processors have market power and want to increase their margin, then the pass-through rate should be accelerated by higher degrees of market power. That is, $\varphi_2^{(I)}$ in equation (3.14) is expected to be positive. Similarly, in regime (*III*) ($\hat{\varepsilon}_{t-d} \geq \gamma_2$), retail prices are relatively high compared to the equilibrium. In this regime, coffee processors with market power may want to keep their margins high by slowing the pass-through rate. Therefore, $\varphi_2^{(III)}$ is expected to be negative, reducing the speed of adjustment to the equilibrium. Regime (*II*) represents deviations of magnitude between the threshold values [γ_1, γ_2]. This regime accounts for the potential presence of price friction underlying price transmission behaviors. As Balke and Fomby (1997) pointed out, retail prices may be expected to respond when the benefits from doing so exceed the costs. The regime between γ_1 and γ_2 in equation (3.13) can be interpreted as the sector where retail prices do not respond to changes in international coffee bean prices because the magnitude of the change is too small to induce price changes at retail (Goodwin & Piggott 2001; Meyer 2004; Meyer & von Cramon-Taubadel 2004). Therefore, the pass-through rate parameters, $\varphi_1^{(II)}$ and $\varphi_2^{(II)}$ are expected to be close to zero and statistically insignificant.

Equation (3.13) can be extended by accommodating short-run asymmetric retail price responses for rising and falling international prices (Wolffram 1971; Houck 1982; von Cramon-Taubadel and Loy 1996). An asymmetric formulation of equation (3.13) is specified by splitting the short-run dynamics variables $\Delta w_{t(t-1)}$ into their positive, $\Delta^+ w_{t(t-1)}$ and negative, $\Delta^- w_{t(t-1)}$ components. Thus, $\Delta^+ w_{t(t-1)} = \Delta w_{t(t-1)}$ if $\Delta w_{t(t-1)} \geq 0$, zero otherwise; and $\Delta^- w_{t(t-1)} = \Delta w_{t(t-1)}$ if $\Delta w_{t(t-1)} < 0$, zero otherwise. Therefore, an asymmetric formulation incorporating regime-specific and market power-dependent pass-through rates is given by:

$$\Delta p_t = \begin{cases} \lambda_0 + \lambda_1^{(I)}\hat{\varepsilon}_{t-1} + \lambda_2 \Delta p_{t-1} + \lambda_3^+ \Delta^+ w_t + \lambda_3^- \Delta^- w_t + \lambda_4^+ \Delta^+ w_{t-1} + \lambda_4^- \Delta^- w_{t-1} + u_t, & if\ \hat{\varepsilon}_{t-d} < \gamma_1 \\ \lambda_0 + \lambda_1^{(II)}\hat{\varepsilon}_{t-1} + \lambda_2 \Delta p_{t-1} + \lambda_3^+ \Delta^+ w_t + \lambda_3^- \Delta^- w_t + \lambda_4^+ \Delta^+ w_{t-1} + \lambda_4^- \Delta^- w_{t-1} + u_t, & if\ \gamma_1 \leq \hat{\varepsilon}_{t-d} < \gamma_2 \\ \lambda_0 + \lambda_1^{(III)}\hat{\varepsilon}_{t-1} + \lambda_2 \Delta p_{t-1} + \lambda_3^+ \Delta^+ w_t + \lambda_3^- \Delta^- w_t + \lambda_4^+ \Delta^+ w_{t-1} + \lambda_4^- \Delta^- w_{t-1} + u_t, & if\ \hat{\varepsilon}_{t-d} \geq \gamma_2 \end{cases}, \quad (3.15)$$

where $\lambda_1^{(k)}(L_t) = \varphi_1^{(k)} - \varphi_2^{(k)} L_t$, $\forall k = I, II, III$.

3.4 Data

The empirical analysis is based on monthly data for the roasted coffee market in the United States and Germany over the period January 1992 to December 2009. We compile coffee price data ($/lb) including retail prices in the United States and Germany, international coffee bean prices of the three varieties under consideration and international composite prices. The latter are the weighted average prices of different coffee bean varieties constructed from the International Coffee Organization (ICO). Coffee consumption for each country is measured as imports minus re-exports (Bettendorf and Verboven 2000). These data are obtained from the ICO. We employ disposable income per capita as our measure of income. The income data for both countries are denoted in U.S. dollars and are obtained from the Bureau of Labor Statistics (BLS 2011) and the Federal Statistical Office (FSO 2011) for the United States and Germany, respectively. We employ a tea price index for the U.S. model and a carbonate soft drink price index for the German model as substitute goods, respectively. These data are also compiled from the Bureau of Labor Statistics and the Federal Statistical Office for the United States and Germany, respectively. The unit labor cost index and the fuel price index are used as proxies for labor and distribution costs, respectively. All prices and income variables are adjusted by Consumer Price Index to take as real values.

Table 3.2 presents descriptive statistics for these data. The descriptive statistics reveal that the average retail price during the sample period is higher in Germany ($4.13/lb) than in the United States ($3.33/lb). Table 3.2 also shows that the degree of variation measured by the coefficient of variation.[8] The results indicate that variability of international bean prices and international composite prices are higher than those of retail prices in both countries. These figures imply that international prices of coffee beans are more volatile than retail prices. On the other hand, coffee bean prices for three varieties seems to be similar in terms of average and their standard deviations for both countries. As described in Table 3.1, total consumption of roasted

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[8] The coefficient of variation is measured as the ratio of standard deviation to mean. This measure is also referred as a relative standard deviation.

coffee is about two times larger in the United States than those in Germany. However, per capita consumption is larger in Germany and that of the United States. The coefficient of variation for consumption shows that variability of consumption seems to be higher in Germany than in the United States, implying a relatively more sensitiveness in the German coffee consumptions.

Table 3.2 Descriptive statistics

	United States		Germany	
	Mean	C.V.	Mean	C.V.
Retail Price ($/lb)	3.33	0.16	4.21	0.22
Other Mild ($/lb)	1.09	0.38	1.13	0.38
Columbian Mild ($/lb)	1.18	0.38	1.17	0.38
Robusta ($/lb)	0.68	0.47	0.67	0.48
Composite price ($/lb)	0.90	0.39	0.90	0.39
Consumption (Million Bushel)	1.62	0.17	0.80	0.29
Per capita disposable income (1000 $)	26.80	0.21	7.79	0.18
Tea/Coke (Price Index) [a]	1.32	0.10	0.98	0.04
Labor (Price Index) [b]	0.96	0.07	0.90	0.10
Distribution (Price Index)[a]	1.11	0.59	0.78	0.29

a. For the U.S. 1982-84 =100 and for Germany 2005 = 100.

b. 2005 = 100

3.5 Empirical results

3.5.1 Demand and supply estimates of roasted coffee

We estimate the simultaneous system shown in equations (3.8) and (3.9) using the iterative GMM. The estimation considers heteroscedasticity and serial correlation of the residuals, using distance matrices for weighting. The distance matrices for the orthogonality conditions are updated at each stage of the iterative process until the vector of coefficients converges. Thus, the regression standard errors and covariance matrices are corrected both for heteroscedasticity and autocorrelation (Ogaki 1999).

For the iterative GMM estimation, we employ a number of instrument variables[9] for model identification, and test whether or not this model is over-identified employing Hansen (1982)'s *J*-statistic. Our over-identification test results show that the null hypothesis of a valid model specification cannot be rejected at 5% significant level. This result suggests that both the U.S. and the German models are correctly specified.

Table 3.3 reports the GMM parameter estimates for the U.S. and the German roasted coffee markets. Overall, our parameter estimates exhibit the expected signs and significances. In the demand equation, the estimated coefficient on price is negative and significant for both countries. Specifically, the estimate coefficients suggest that a $1 increase in retail prices in the U.S. and the German market leads to 0.32 million bushel and 0.09 million bushel decreases in

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[9] We use three different international bean prices including *Other Mild*, *Columbian Mild* and *Robusta*, price indices for labor and distribution, household expenditure, exchange rate for Germany and import price index for the United States.

total quantities of roasted coffee consumed, respectively.[10] This parameter estimate, $\hat{\alpha}_1$, is employed to derive the price elasticity of demand. The coefficients on the income variables for both countries are also positive and significant, as expected. The results show that when per capita disposable income increases by $1, total roasted coffee consumption increases by 82 thousand and 87 thousand bushels in the United States and Germany, respectively. Given differences in total population, German coffee consumers seem to be more responsive to changes in income than those in the United States. The time trend variable shows a gradual decrease in roasted coffee consumption in both countries. However, the magnitude to which per capita consumption decreases is slightly larger in the German market than in the U.S. market. The coefficient on tea prices in the U.S. model is negative and significant. These results indicate that higher tea prices lead to lower coffee demand, which is contrary to our expectation. In the German model, the results suggest that there is no significant effect of the price of substitutes on coffee consumption.[11]

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[10] These values correspond to 0.0001737g/capita and 0.0001793g/capita decreases for the U.S. and the German market, respectively, indicating that German consumers seem to be slightly more sensitive to changes in retail prices than U.S. consumers.

[11] Feuerstein (2002) argues that there is no substitute good for roasted coffee in Germany, whereas Kutty (2000) and Bettendorf and Verboven (2000) consider tea as a close substitute.

Table 3.3 Roasted coffee demand and supply estimates in the U.S. and the German markets

	U.S.		Germany	
	Coefficient	Std. Errors	Coefficient	Std. Errors
Demand				
Constant	2.915**[a)]	(0.636)[b)]	0.482	(0.670)
Price	-0.324**	(0.084)	-0.088**	(0.023)
Time	-0.008**	(0.003)	-0.003**	(0.001)
Income	0.082**	(0.026)	0.087**	(0.017)
Tea/Coke	-2.677**	(0.584)	0.307	(0.616)
Supply				
Constant	-2.241**	(0.906)	-3.737**	(4.181)
Other Mild	-1.239	(0.662)	1.175**	(0.437)
Columbian Mild	2.244**	(0.616)	-0.640	(0.456)
Robusta	0.512**	(0.255)	2.655**	(0.299)
Labor	3.630**	(1.050)	0.040**	(0.013)
Distribution	-0.222**	(0.110)	-1.975**	(0.463)
Conjectural Variation Elasticity (θ)	0.268**	(0.127)	0.417**	(0.164)
Price Elasticity (η)[d]	0.403		0.547	
Lerner Index (*L*)[d]	0.665		0.762	
Hansen's *J* Statistics (Over-identifying tests)	8.13	(0.087)[c)]	18.59	(0.100)
Number of Observations	216		216	

a. ** and * indicate 5% and 10% significant level, respectively.

b. Parenthesis indicates standard errors.

c. Parenthesis indicates *P*-value.

d. Price elasticities and the Lerner Index are evaluated at sample means.

As in equation (3.9), retail prices of roasted coffee are determined by two factors: the first factor represents the impacts of marginal costs. Recall that our model considers coffee processors and retailers as a single firm. Therefore, marginal costs in this model may include roasting costs of coffee beans for processors and distribution costs of roasted coffee for retailers; the second factor refers to the impacts of the extent of oligopoly power in retail pricing.

We first interpret the estimated parameters on the marginal cost measure. Given the relatively fixed proportion of other input factors such as labor and distribution in determining marginal costs, coffee bean prices have a substantial effect on the variation of the industry level marginal costs. In coffee processing, *Other Mild* and *Columbian Mild* are the primary varieties used in roasted coffee production and their mix ratio is closely related to product quality. *Columbian Mild* variety is employed in the production of premium roasted coffee. On the other hand, *Robusta* variety is only employed as an additive (in small amounts) in the production of roasted coffee. These varieties are used mostly in the production of instant coffee. In this regard, our parameter estimates show substantive differences in the structure of marginal costs between the two countries. The coefficients for *Columbian Mild* and *Robusta* prices have a positive and significant effect on marginal costs in the U.S. model. That is, marginal costs of the U.S. roasted coffee industry appear to be directly affected by variations in prices of *Columbian Mild* and *Robusta* varieties. In contrast, changes in *Other Mild* variety prices have a modest impact on marginal costs in the U.S. These results imply that the *Columbian Mild* variety is preferred to the *Other Mild* variety in the United States. Specifically, a 1$ increase of *Columbian Mild* and *Robusta* prices leads to a $2.24 and a $0.51 increase of marginal costs, respectively. The higher estimate of the coefficient on *Columbian Mild* prices may be explained by the fact that our model does not include marketing costs, which are presumably larger for premium products.

The results are different from the German market. The marginal costs in the German roasted coffee sector seem to be driven by prices of *Other Mild* and *Robusta* varieties. These results imply that these two varieties, *Other Mild* and *Robusta* are the primary input in the production of roasted coffee in this country, reflecting the German consumers' preferences in taste. Similarly, higher estimated values of coefficients on *Other Mild* and *Robusta* also seem to take over marketing costs which are not reflected in our data.

We now focus on the interpretation of the market conduct parameters, which are closely related to pricing behaviors in an oligopolistic market environment. Pricing decisions in a homogeneous product oligopolistic market are influenced by the price elasticity of demand, the conjectural variation elasticity, and the marginal costs (Iwata 1974). Given that the marginal cost structure for roasted coffee processing is relatively similar,[12] disparities in pricing behavior depend on differences in the consumers' sensitivity to price changes in the demand side and the degree of competitiveness of firms on the supply side. First, we derive the absolute value of the price elasticity of demand evaluated at the sample mean. The price elasticities of demand are estimated at 0.403 and 0.547 for the United States and Germany, respectively. These results imply that German consumers are more sensitive to retail price changes of roasted coffee than U.S. consumers are. Second, the estimated conjectural variation elasticities for both countries are significant and range between 0 and 1, as expected. From equation (3.9), an elasticity of conjectural variation greater than zero implies that retail prices are set a higher level than the marginal costs. Our results thus provide evidence of non-competitive pricing behavior in both

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[12] In the cost structure of roasted coffee processing, the largest share of costs is green coffee beans (about 60 percent of total processing costs) for which the U.S. and Germany roasters purchase raw materials at world market prices.

countries, further indicating that the roasted coffee industries in the United States and Germany are oligopolistic.

These two measures are employed to derive the Lerner Index which measures the degree of market power, as in equation (3.5). The Lerner Index indicates to what extent the retail prices are set higher than the marginal costs. Our results suggest that the roasted coffee industries in both countries exert considerable market power in retail pricing. Specifically, the Lerner Index of the German roasted coffee market calculated at sample mean is 0.762, which is greater than the Lerner Index of the United States, calculated at 0.665. The difference is mainly driven by relatively higher conjectural variation elasticity in Germany. This implies that the German coffee industry seems to exercise more market power than its U.S. counterpart. This result supports the observation that the supply sector in the German roasted coffee market is more concentrated than in the United States (see Table 3.1).

3.5.2 Cost pass-through

We now use our estimates of the time-varying Lerner Index to evaluate whether market power is linked to cost pass-through behaviors of the roasted coffee market in the United States and Germany. The time-varying Lerner Index are calculated using the estimated conjectural variation elasticity and time-dependent price elasticity of demand, as in equation (3.5). To analyze cost pass-through behaviors in these two countries, our strategy is as follows: First, we examine the time series properties (e.g. integration and co-integration) of two prices series, international (composite) prices, w_t, and retail prices in both countries, p_t. Second, we estimate the TAR model in equation (3.10) to identify the unknown threshold values and to test their statistical significance. Lastly, we find empirical evidence on impacts of market power on the cost pass-through rates in these two countries by estimating the TECM incorporating the market power measures as in equation (3.15).

3.5.2.1 Integration and Co-integration

In this section, we examine the time series properties of price variables including stationarity of two price series, w_t and p_t, and co-integration between them. To examine integration, we employ the augmented Dickey-Fuller (ADF) test under the null hypothesis of nonstationarity and the KPSS test under the null hypothesis of stationary. Panel A of Table 3.4 reports both test statistics for retail prices in the U.S. and Germany and for international prices. The ADF test results suggest that the null hypothesis of nonstationarity cannot be rejected in all variables in levels. In contrast, the test for the variables in first differences rejects the null hypothesis, indicating that they follow *I(0)* processes. The KPSS test results support these findings. The KPSS test indicates that all price variables in levels exhibit a unit-root, but all variables in first differences follow a stationary process.

Another issue in the ECM approach is whether the two price series have a long-run linear relationship. We test the co-movements between international and retail prices employing Johansen's (1992a, 1992b, 1995) approach. We conduct *λmax* and *trace* tests between retail and international prices in both countries. Panel B of Table 3.4 indicates that, for both countries, there exist at least one co-integration rank, indicating the presence of a co-integration relationship between the two price series in both countries.

Table 3.4 Integration and co-integration tests

A. Integration test

Variables in Levels			U.S. Retail Price	Germany Retail Price	Coffee Bean Price
ADF-t	$H_0 : \sim I(1)$		-2.77	-1.49	-2.51
	$H_0 : \sim I(1)$	*no constant*	-0.11	-0.13	-0.16
KPSS	$H_0 : \sim I(0)$	*no constant*	0.36**	0.64**	0.46**
	$H_0 : \sim I(0)$	*no linear trend*	0.36**	0.82**	0.45*

Variables in First Differences			Δ U.S. Retail Price	Δ Germany Retail Price	Δ Coffee Bean Price
ADF-t	$H_0 : \sim I(1)$		-9.70**	-10.36**	-12.51**
	$H_0 : \sim I(1)$	*no constant*	-9.72**	-10.38**	-12.52**
KPSS	$H_0 : \sim I(0)$	*no constant*	0.05	0.11	0.07
	$H_0 : \sim I(0)$	*no linear trend*	0.06	0.17	0.07

B. Co-integration test

	H_0: r	Critical Value	U.S. Retail Price	Germany Retail Price
$\lambda_{\max}$	0	11.44	34.98	18.73
	1	3.84	0.00	0.04
trace	0	12.53	34.98	18.77
	1	3.84	0.00	0.04

3.5.2.2 The TAR estimate results

Given that each price series is integrated of order one, *I(1)*, and both are co-integrated, we estimate the TAR model in equation (3.10). Table 3.5 presents estimation results of the TAR model.

Table 3.5 Threshold autoregressive (TAR) estimates

	U.S.	Germany
Optimal Lag	7	5
Delay Parameter	1	1
Tsay (1997) Test	3.34** (0.00) [a]	7.51**(0.00)
Estimated Threshold		
γ_1	-0.1896	-0.6997
γ_2	-0.3357	-0.7537
LR statistics	5.41	13.38
Bootstrapped *p* value	(0.00)[b]	(0.00)
Percentage of observations by regime		
Regime I	0.13	0.09
Regime II	0.09	0.01
Regime III	0.78	0.90

a. *F*-test for a linear AR process and parenthesis shows asymptotic *p* values for test statistics.

b. 300 bootstrap replications were used to obtain the *p* value.

We choose the optimal lags of 5 and 7 for the U.S. and German model, respectively, using the AIC and SBC measures. The delay parameter, *d* is selected to 1 for both countries based on the *TAR-F* statistics in *Tsay* test (Goodwin and Hold 1999; Goodwin and Piggott 2011). The *Tsay* (1989) test results suggest that the null hypothesis of no threshold effects in the autoregressive process is rejected at the 5% significant level, indicating the existence of nonlinear threshold effects in the error correction process of variable $\hat{\varepsilon}_t$.

Given the presence of threshold effects, we estimate the unknown threshold parameters, γ_1 and γ_2, following Chan (1983) and Hansen (1999). The estimated threshold parameters are -0.336 and -0.189 for the U.S. model; and are -0.754 and -0.699 for the German model. The *LR* statistics and their bootstrapped *p*-values indicate that the estimated multiple threshold values are statistically significant. The percentage of observations falling in regimes *I*, *II*, and *III* is 13, 9 and 78 percent in the United States and 9, 1 and 90 percent in Germany, respectively.

The estimated threshold values and distributions of observations falling in each regime reveal important preliminary facts associated with the cost pass-through behaviors in the two countries. Recall that the error correction terms derived from the co-integration relationship between retail and international prices are the deviations from the long-run equilibrium relationship. A positive deviation means that retail prices are high relative to the long-run equilibrium level, implying that international prices are falling. In contrast, a negative deviation indicates that retail prices are below the long-run equilibrium level, implying that international prices are rising. The results indicate that the threshold values partitioning regime *I* for the United State and Germany are estimated at -0.336 and -0.754, respectively. That is, retail prices are low relative to the long-run equilibrium level. Therefore, distribution in importing countries is expected to increase retail prices to restore the equilibrium. Interestingly, the estimated threshold values partitioning regime *III* are -0.189 and -0.699 for the United States and Germany, respectively. Therefore, this regime includes all positive values of the error correction term (i.e. when retail prices are too high relative to the long run equilibrium), but also some values in which this term is negative (e.g. [-0.189, 0] for the U.S.; [-0.669, 0] for Germany). In other words, for modest increases in international prices (i.e. retail prices are relatively little low compared to the equilibrium), the distribution system adjusts toward the long run equilibrium in the same manner as when the retail prices are high. The percentage of observations falling in between [γ_2, 0] of regime *III* observations for the U.S. and the German models is 29 percent and 46 percent, respectively.

3.5.2.3 Empirical results of cost pass-through

In this section, we investigate cost pass-through behaviors reflecting the impact of market power in the United States and Germany, employing the TECM described in equations (3.14) and (3.15). Our analysis consists of the two steps. In the first step, we examine the existence of short-run asymmetries in price transmission. In the second step, we interpret our estimated pass-through rates and the influence of market power on them.

To examine the existence of short-run asymmetries, we test the null hypothesis of equality between the estimated parameters on variables in first differences for positive and negative changes in international prices (i.e. $\lambda_{3(4)}^{+} = \lambda_{3(4)}^{-}$). We present these results in Table 3.6. Our findings provide evidence of short-run asymmetries in parameters on the lagged variable in the United States, but not in Germany. We also employ the AIC and the SBC to explore whether or not an asymmetric model is appropriate to examine the relationship between international and retail prices. The calculated AIC and SBC values indicate that an asymmetric model is preferable to a symmetric model in the United States. In contrast, a symmetric TECM is more appropriate than its asymmetric counterpart in Germany (Table 3.6).

Table 3.6 Model selection criteria

		U.S.	Germany
F-Statistic (H_0: $\lambda_3^+ = \lambda_3^-$)		1.88 (0.17)	1.95 (0.16)
F-Statistic (H_0: $\lambda_4^+ = \lambda_4^-$)		25.26** (0.00)	0.03 (0.86)
Symmetric	*AIC*	142.47	274.99
TECM	*SBC*	176.13	308.02
Asymmetric	*AIC*	110.95	277.76
TECM	*SBC*	154.71	321.52

a. Parenthesis indicates the *p* value.

Second, given the model selection results, we analyze the cost pass-through behaviors in the U.S. and the German roasted coffee markets. These results are presented in Table 3.7. Recall that the pass-through rate is defined as the extent to which retail prices adjust to reestablish the long run equilibrium after a change in international coffee bean prices. A measure of this adjustment is provided by the parameter estimates of the coefficient on the error correction terms $\lambda_1^{(k)}$, where *k*=*I*, *II* and *III*.

Table 3.7 Threshold error correction model (TECM) estimates

	U.S.		Germany	
Constant	-0.046	(0.011)[b]	0.002	(0.011)
$ECT_{t-1}^{(I)}$	-0.296**[a]	(0.101)	-0.257**	(0.086)
$L_t \cdot ECT_{t-1}^{(I)}$	0.260*	(0.144)	0.152*	(0.081)
$ECT_{t-1}^{(II)}$	0.514	(0.399)	-0.035	(0.246)
$L_t \cdot ECT_{t-1}^{(II)}$	-0.660	(0.466)	0.083	(0.220)
$ECT_{t-1}^{(III)}$	-0.231**	(0.095)	-0.098	(0.063)
$L_t \cdot ECT_{t-1}^{(III)}$	0.219	(0.143)	0.054	(0.069)
Δp_{t-1}	0.251**	(0.054)	0.289**	(0.064)
Δw_t	--		0.336**	(0.105)
$\Delta^+ w_t$	0.111	(0.109)	--	
$\Delta^- w_t$	-0.190	(0.166)	--	
Δw_{t-1}	--		-0.098	(0.114)
$\Delta^+ w_{t-1}$	0.876**	(0.137)	--	
$\Delta^- w_{t-1}$	-0.357**	(0.166)	--	
$\varphi_1^{(I)} - \varphi_1^{(I)} \cdot L_t$	-0.470		-0.653	
$\varphi_1^{(II)} - \varphi_1^{(II)} \cdot L_t$	--		--	
$\varphi_1^{(III)} - \varphi_1^{(III)} \cdot L_t$	-0.231		--	
DW	2.12		1.93	
R^2	0.55		0.24	

b. ** and * indicate 5% and 10% significant level, respectively.

c. Parenthesis indicates standard errors.

We discuss the results for the United States first. For this country, the estimated parameters on the pass-through rates($\varphi_1^{(k)}$) which is independent of market power measure (L_t) are negative and significant for regimes *I* and *III* as predicted by theory. In contrast, this coefficient is not significant in regime *II*. This is consistent with argument that there may be a "no-adjustment range" due to the presence of price rigidities, as reflected by the price transmission behavior observed in regime *II*. The estimated pass-through rates for regime *I* and *III* are -0.296 and -0.231, respectively. These results indicate that the magnitude of the pass-through rate is slightly larger in regime *I* than regime *III* in the United States. Therefore, changes in international prices tend to be passed on to consumers slightly faster when international prices are rising than when they are falling. The pass-through rate could be also affected by market power as we hypothesize in equation (3.14). This is measured by the parameter estimates $\varphi_2^{(k)}$, for *k*=*I*, *II* and *III*. Our results show fundamental differences between regimes. That is, market power has a significant effect on the pass-through rate in regime *I*, whereas it has no influence on the pass-through rate in regime *III*. These results indicate that when international prices are increasing, retail prices tend to increase faster by the exertion of market power. However, when international prices are decreasing, we find no additional effect by the exertion of market power. The overall pass-through rates, λ_1^I and λ_1^{III} including the market power effects are estimated at -0.470 and -0.231 for regime *I* and *III*, respectively.[13] These results provide evidence of "rockets and feathers" behavior. That is, market power seems to cause retail prices to increase quickly and decrease slowly in the U.S. coffee supply chain.

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[13] These values are computed by $\hat{\varphi}_1^{(k)} - \hat{\varphi}_2^{(k)}\overline{L}_t$ where *k*=*I* and *III*.

Our results for the German market are also reported in Table 3.7. We analyze the estimates results of the symmetric TECM, following the results from Table 3.6. Similar to the U.S. results, the parameter estimates on regime *II* are not significant, which supports the "no adjustment range" argument associated with the presence of price rigidities. However, our results show differences in pass-through behaviors between the United States and Germany. The estimated pass-through parameters ignoring the influence of market power indicate that the estimated pass-through rate in regime *I* is -0.257, but statistically insignificant in regime *III*. These results imply that the pass-through occurs only when international price are rising. This is indicative of "rockets and feathers" behavior in the German market, regardless of the influence of market power. Moreover, our results show that market power seems to accelerate the response of retail prices in regime *I*, whereas it has no impact on the pass-through rate in regime *III*. The overall pass-through rate incorporating the effect of the market power is -0.653 in regime *I*; whereas in regime *III* there is no significant pass-through behavior for a negative changes in international prices. These results imply that, in the German coffee industry, both market power and other factors seem to lead to the "rockets and feathers" phenomenon in retail pricing.

3.5.2.4 Discussion

Our findings on the market conduct parameters and the cost pass-through behaviors suggest that market power influences the pass-through rate only in regime *I*. These results indicate that when retail prices are below the equilibrium, the pass-through rate becomes faster by the exertion of market power. This is perhaps because the industry wants to avoid falling profits by quickly passing increases in coffee bean prices on to consumers (Ward 1982). In contrast, when retail prices are above the equilibrium, the industry seems to maintain existing retail prices until demand conditions force a change (Peltzman 2000).

However, there are some differences in retail pricing behavior between the U.S. and the German roasted coffee industries. One difference is observed in terms of the magnitude of market power effects on the pass-through rates. The extent to which the pass-through rate is increased by the market power is higher in Germany than in the United States. In Germany, market power raises the pass-through rate by about 150 percent, whereas it leads to about 60 percent faster rise of retail prices in the United States. These results are in part due to differences in market conduct given that the degree of oligopoly power measured by the Lerner Index is stronger in Germany than in the United States. Another difference between the two countries is the underlying ground of the "rockets and feathers" behaviors in the cost pass-through. In the U.S. roasted coffee industry, the exertion of market power seems to be a crucial factor causing the "rockets and feathers" phenomenon. Without the effect of market power on cost pass-through rates, the estimated coefficients of cost pass-through variable on regime *I* and *II* are similar (e.g. -0.296 for regime *I*; -0.231 for regime *III*). That is, retail prices seem to be accelerated mostly by the exertion of market power in the United States. In contrast, in the German roasted coffee industry, the estimated coefficient of cost pass-through variable, excluding the effect of market power, is -0.257 for regime *I*, but is insignificant for regime *III*. This indicates that, without the effect of market power, there is evidence on asymmetric cost pass-through rates showing the "rockets and feathers" principles. That is, in the German roasted coffee market, although the exertion of market power causes retail price respond faster for increases in international prices than for decreases, other factors different than market power also lead to this conventional phenomenon.

3.6 Concluding remarks

In this study, we examine the links between non-competitive pricing behaviors and cost pass-through patterns in the roasted coffee markets of the United States and Germany, the two largest coffee importing countries. Using aggregate monthly data from January 1992 to December 2009, we first evaluate pricing behaviors by estimating the price elasticities of demand and the conjectural variation elasticities in the two countries. Subsequently, we use these two measures to obtain a Lerner Index, and thereby to evaluate the degree of market power of the roasted coffee industry in the two countries. Second, we estimate the cost pass-through rates that incorporate the influence of market power.

We find that the roasted coffee markets in both countries are characterized by oligopolistic structures. The degree of market power is stronger in Germany than in the U.S., driven by the relatively higher conjectural variation elasticity in Germany. This result supports the market observation that the coffee supply chain in Germany is more concentrated than in the United States. Using our market conduct measures, we find evidence that market power leads to an increase of the pass-through rates when retail prices are below equilibrium. These results provide evidence of "rockets and feathers" behavior in retail pricing in both countries. The "rockets and feathers" phenomenon seems to be mainly caused by the exertion of market power in the United States. In contrast, other factors different than market power seem to also lead to this conventional phenomenon in Germany.

Our study makes valuable contributions to the literature by providing policy makers concerned with the performances of food supply chain with meaningful empirical tools for identifying underlying causes associated with imperfect cost pass-through. But our approach has some limitations that should be addressed in future research. First, our study aggregates coffee roasters and retailers in importing countries. Future studies should conduct firm-specific analyses and explore the strategic pricing behavior between the coffee processing and retailing sectors. Second, the homogeneity assumption of the good 'roasted coffee' is somehow restrictive. The national markets in both countries are characterized by extensive product differentiation with multiple brands and blends offering different qualities. Future research should thus develop models incorporating production differentiation.

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CHAPTER 4

DISSERTATION CONCLUSION

This dissertation studied three topics rooted in complex vertical and spatial relationships existing in supply chains for specialty crops: 1) impact of policies to control CO2 emissions on spatially- and temporally-disaggregated food supply chains; 2) vertical price transmission changes after elimination of an export cartel; and 3) the influence of market power on cost pass-through.

Chapter 1 investigated the impact of two alternative CO2 emission reduction policies, namely a carbon tax and a cap-and-trade system, on the economic performance of supply chain for fresh apples. The Chapter also considered the potential impact of incentivizing farm-level CO2 sequestration options. A spatially- and temporally-disaggregated price equilibrium model is constructed to evaluate potential tradeoffs between CO2 emission reduction initiatives and economic performance. The results show that all CO2 emission policies examined lead to modest reductions in CO2 emissions along the fresh apple supply chain when carbon price ranges between \$25 and \$200 per metric ton of CO2. The results also suggest that a cap-and-trade system is relatively more cost-effective for CO2 emission reduction than a carbon tax, regardless of the presence of CO2 sequestration options. The models and simulation results in this Chapter provide useful insights to decision makers promoting industry- or sector-specific initiatives intended to reduce carbon footprints. Furthermore, the model developed here can assist policy-makers in the identification of socially acceptable carbon prices and tax levels that also satisfy both CO2 emission reductions and social welfare losses through the simulation of various policy scenarios.

Chapter 2 examined how the end of the coffee export quota system (EQS) influenced international-to-retail price transmission behaviors in supply chains for roasted coffee in France, Germany and the United States. A threshold error correction model (TECM) was developed to measure price transmission behaviors, taking into account long-run threshold effects and short-run price transmission asymmetries. The empirical findings suggest that the elimination of the EQS brought about fundamental differences in price transmission behaviors in each of these countries. Specifically, the results indicate that retail prices became more responsive to changes in international prices after the end of the EQS. Moreover, the results show differences in short-run price transmission behaviors across countries in the post-EQS period. These differences were discussed in terms of country-specific market structures. The TECM employed in this Chapter improves the accuracy of impact assessments of policy changes on price transmission processes compared to approaches employed in earlier. Therefore, the methodology in this Chapter provides meaningful tools for private and public decision-makers who are concerned with policies that influence prices in the context of international trade for agricultural commodities.

Chapter 3 measured the influence of market power on cost pass-through rates (i.e. the extent to which changes in international prices of coffee beans are passed on to the end consumers of roasted coffee) in the U.S. and the German roasted coffee market. A structural supply and demand model was developed to establish a time-varying Lerner Index. The index measures the degree of market power, based on the parameter estimates of the structural model. Subsequently, a TECM was used to test the impacts of market power on cost pass-through behaviors in these countries. We find evidence that market power leads to an increase of the pass-through rates when retail prices are below the equilibrium price. These results provide evidence of "rockets and feathers" behavior in retail pricing in both countries. The "rockets and feathers" phenomenon seems to be mainly caused by the exertion of market power in the United States. In contrast, other factors different than market power seem to also lead to this conventional phenomenon in Germany. The approach employed in Chapter 3 remedies some of the shortcomings of the methods used in Chapter 2, given that the empirical findings in Chapter 2 were silent regarding the factors that trigger non-competitive price transmission behaviors. Therefore, the models and results presented in Chapter 3 provide private and public decision-makers who are interested in monitoring anti-trust conduct in food supply chains with meaningful tools to identify anti-competitive pass-through behaviors.

Overall, this dissertation provides a toolkit of quantitative models (nonlinear optimization models and time series econometrics) that foster a better understanding of critical issues influencing the performance of supply chains for specialty crops. Such issues include the impact of various policy interventions (CO2 emission reduction initiatives and trade liberalization) on prices and economic performance; and the implications of market power exertion on cost pass-through. These models can be applied to other specialty crop supply chains and provide valuable information to private and public decision-makers in the design of interventions affecting the performance of food supply chains.