

**Essays on Firm Heterogeneity with Empirical
Applications in Economic History and Agricultural
Economics**

By

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Abstract

This thesis uses models of firm-heterogeneity to complete empirical analyses in economic history and agricultural economics. In Chapter 2, a theoretical model of firm heterogeneity is used to derive a statistic that summarizes the welfare gains from the introduction of a new technology. The empirical application considers the use of mechanical steam power in the Canadian manufacturing sector during the late nineteenth century. I exploit exogenous variation in geography to estimate several parameters of the model. My results indicate that the use of steam power resulted in a 15.1 percent increase in firm-level productivity and a 3.0-5.2 percent increase in aggregate welfare. Chapter 3 considers various policy alternatives to price ceiling legislation in the market for production quotas in the dairy farming sector in Quebec. I develop a dynamic model of the demand for quotas with farmers that are heterogeneous in their marginal cost of milk production. The econometric analysis uses farm-level data and estimates a parameter of the theoretical model that is required for the counterfactual experiments. The results indicate that the price of quotas could be reduced to the ceiling price through a 4.16 percent expansion of the aggregate supply of quotas,

or through moderate trade liberalization of Canadian dairy products. In Chapter 4, I study the relationship between farm-level productivity and participation in the Commercial Export Milk (CEM) program. I use a difference-in-difference research design with inverse propensity weights to test for causality between participation in the CEM program and total factor productivity (TFP). I find a positive correlation between participation in the CEM program and TFP, however I find no statistically significant evidence that the CEM program affected TFP.

Dedication

To my parents, Farrar, and my children.

Acknowledgment

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Chapter 1: Introduction

In recent years theoretical and empirical research on firm-level heterogeneity has made important contributions to many fields of economics, notably including international trade and industrial organization. This literature has focused on understanding how the allocation of production across firms with varying levels of productivity affects aggregate outcomes. For example, the theoretical frameworks of Hopenhayn (1992) and Melitz (2003) have been used extensively to study the effects of economic policy on the distribution of firm productivity and market share. Empirically, the increasing availability of firm-level data has been accompanied by innovations in econometric theory, such as the development of new methodologies for estimating total factor productivity at the firm-level. The objective of this thesis is to apply theoretical and empirical models of firm-heterogeneity in two sub-fields of economics where these frameworks have less commonly been employed: economic history and agricultural economics.

In Chapter 2, I use a model of firm heterogeneity to estimate the effect of mechanical steam power on firm-level productivity and aggregate welfare. The empirical

application uses complete count firm-level data from the 1871 Canadian Census of Manufacturing. This historical setting provides an ideal environment to study the relationship between industrialization, technological adoption, and the distribution of firm productivity and scale. I develop an econometric framework that addresses the inherently endogenous relationship between technological adoption and unobserved firm-level productivity using instrumental variables estimation. The econometric estimates are used to calculate the aggregate gains from steam power using a welfare statistic that is derived from the theoretical model. My results indicate that the use of steam power resulted in a 15.1 percent increase in firm-level productivity and a 3.0-5.2 percent increase in welfare. I conclude that using a model that explicitly accounts for firm-heterogeneity in adoption results in qualitatively larger estimates of the gains from steam power, as compared to previous estimates that have used a growth accounting framework. The analysis in Chapter 2 also contributes to the literature by providing a theoretical and empirical framework that can be used to study the firm-level productivity and aggregate welfare gains from the adoption of other technologies during other eras.

In Chapter 3, I consider various policy alternatives to price ceiling legislation in the market for production quotas in the dairy farming sector in Quebec. The system of supply management in the Canadian dairy sector requires that farmers hold quotas to produce milk. In Canada's two largest dairy producing provinces, Quebec and Ontario, ceilings on the price of quotas have been in effect since 2007

and 2009 respectively. Previous research established that the use of quota price ceilings create a source of inefficiency in the Canadian dairy sector. I contribute to the literature on this topic by developing a dynamic model of demand for quotas with farmers that are heterogeneous in their marginal cost of milk production. I develop an econometric model and use farm-level data to estimate a parameter of the model that is required for the counterfactual policy experiments. Using my econometric results and the modeled equilibrium price, I estimate that in 2011 the value of dairy quotas in Quebec was approximately 20% higher than the ceiling price of \$25,000 per unit. The counterfactual experiments indicate that the price of quotas could be reduced to the ceiling price level through a 4.16% expansion of the aggregate supply of quotas, or through moderate trade liberalization of Canadian dairy products. I estimate that removing the price ceiling and reducing the farm price of milk by 9.97% would result in a 34.6% increase in the minimum viable scale of production.

In Chapter 4, I study the relationship between farm-level productivity and Canadian dairy farmers' participation in the Commercial Export Milk (CEM) program during the period 2000-2003. The CEM program permitted farmers to sell milk directly to Canadian processors provided that the final goods were destined for export. Farmers could sell milk through the CEM program without holding production quotas, however they were paid a farm price that was approximately half of the domestic price under supply management. I find a positive correlation between participation in the CEM program and farm-level total factor productivity. I then use a difference-

in-difference research design with inverse propensity weights to test for causality in the relationship between participation in the CEM program and TFP. I find no statistically significant evidence that the CEM program affected farm-level TFP. To my knowledge, this is the first paper to test for a causal effect of export participation on farm-level productivity in the agricultural sector. Studying the relationship between exporting and productivity at the farm-level is important, as agriculture remains one of the few sectors where substantial trade liberalization has yet to occur. This paper also contributes to the ongoing debate over trade liberalization of Canada's supply-managed agricultural industries within the context of the Canada-EU Comprehensive Economic and Trade Agreement and the Trans-Pacific Partnership.

This thesis is comprised of five chapters. Chapter 2 studies the welfare and firm-level productivity gains from mechanical steam power adoption in the Canadian manufacturing sector in the nineteenth century. Chapter 3 analyzes the implications of price ceiling legislation on production quotas in the Canadian dairy sector. Chapter 4 studies the relationship between the CEM Program and farm-level TFP in the Canadian dairy sector. Chapter 5 concludes.

Chapter 2: Firm Heterogeneity and the Gains from Technological Adoption: Theory and Measurement

2.1 Introduction

There are several theoretical and empirical challenges associated with estimating the aggregate effects of technological adoption. Aggregate models of technological adoption abstract from important questions regarding why only some firms adopt certain technologies and how this heterogeneity affects aggregate outcomes. While models of firm heterogeneity are increasingly popular, tractability in aggregation poses a challenge to estimating the macroeconomic effects of technological adoption in micro-founded models. From an empirical perspective, the inherently endogenous relationship between productivity and technological adoption raises issues of identification that are not easily resolved. This paper develops and estimates a model of firm

heterogeneity and technological adoption that overcomes these challenges.

I develop a model of technological adoption with monopolistically competitive firms that are heterogeneous with respect to their productivity. Building on the framework of Bustos (2011), firms make a discrete choice to adopt one of several different technologies that vary in terms of productivity and their fixed cost of adoption. Despite the superiority of the most productive technology, firm-level heterogeneity and differences in the cost of adoption result in each of the various technologies being used by some fraction of firms in equilibrium. I use the theoretical model to derive a welfare statistic that measures the welfare gains from the introduction of a new technology that is superior to existing technologies. Intuitively, the welfare statistic depends on the counterfactual revenue of firms using the new technology, had they been forced to operate instead with the second most productive technology.

For the empirical application I consider the case of steam power in the late nineteenth century. Studying steam power adoption is a well-suited application of the model for a number of reasons. As a general purpose technology (GPT), my analysis provides insight into the aggregate effects of GPTs when there is firm heterogeneity in adoption.¹ Studying steam-power adoption in Canada is representative of the second generation of countries that entered into industrialization during the late nineteenth century, such as Argentina, Australia, and Sweden. The data used in this paper covers the universe of firms in the Canadian manufacturing sector in 1871. Given the objectives of this paper, the universal nature of the data renders it superior to comparable data for the U.S., such as the Atack and Bateman (1999) samples from the nineteenth century U.S. Manufacturing Censuses. This feature of the data ensures that the aggregate welfare estimates are representative, and provides an adequate sample size for

¹Other papers that have analyzed steam power as a GPT include Crafts (2004) and Rosenberg and Trajtenberg (2004).

consistent estimation of the effect of steam-power adoption on firm-level productivity.

As noted at the onset, an empirical challenge associated with estimating the effect of technological adoption relates to identification. My identification strategy draws on the theory that variation in agricultural land endowments partially explains differential rates of industrialization in the nineteenth century, as has been argued by Rothbarth (1946), Habakkuk (1962), and McCallum (1980). Using an instrumental variables identification strategy, I estimate that steam power adoption increased firm-level labor productivity by approximately 15.1 percent relative to water power. To calculate the welfare statistic, I estimate two structural equations from the theoretical model. My results indicate that steam power adoption resulted in a 3.0-5.2 percent increase in welfare. The empirical framework developed in this paper draws from methodology that has been used to estimate the welfare gains from international trade in models with heterogeneous firms. Comparing my results to this literature, I show that the estimated welfare gains from steam power adoption are approximately one and a half times as large as contemporary estimates of the welfare gains from trade.

This paper contributes to the literature on GPTs by developing a micro-founded empirical framework for estimating the aggregate welfare gains from the introduction of a GPT when firm-heterogeneity factors heavily in the pattern of adoption. My theoretical framework extends Bustos's (2011) model of technological choice by allowing for multiple technologies and by deriving a welfare statistic that can be estimated using firm-level data. This paper also contributes to the literature on steam power, by quantifying the effect of adoption on firm-level productivity and aggregate welfare. There is an extensive literature that documents the importance of steam power to growth and industrialization during the nineteenth century.² My work is similar to

²For the U.S., see Temin (1966), Atack (1979), Atack et al. (1980), Hunter (1985) and Atack et

Crafts (2004), who estimates the aggregate effect of steam power on economic growth in the British economy over the period 1760-1910. Relative to information and communications technologies (ICTs), Crafts (2004) finds that steam technology resulted in only modest contributions to labor productivity growth, peaking during the period 1850-1870 at 0.41 percent per year. However, Crafts (2004) uses a growth accounting framework which ignores the effects of firm heterogeneity. My results suggest that gains stemming from firm heterogeneity in the pattern of steam power adoption were important, as the use of steam power was highly correlated with firm size and productivity.

The remainder of the paper is organized as follows: Section 2.2 describes the data on steam power adoption, Section 2.3 presents the theoretical model, Section 2.4 presents the empirical application of the model to steam power, and Section 2.5 concludes.

2.2 Firm Heterogeneity in GPT Adoption: the Case of Steam Power

An objective of this paper is to develop a tractable framework for estimating the gains from the introduction of a GPT that explicitly accounts for the effects of firm heterogeneity in adoption. The application of this framework in Section 2.4 considers the case of steam power in the Canadian manufacturing sector in the late nineteenth century. This section presents a number of empirical results regarding the use of steam power in Canadian manufacturing during this era that guide the development of the theoretical and empirical frameworks that are used in the remainder of the

al. (2008). For Canada, see Bloomfield and Bloomfield (1989) and Inwood and Keay (2012).

paper.

This paper uses data from the Canadian Industry in 1871 Project (CANIND71), which is a digitized record of the universe of manufacturing establishments that were enumerated in the 1871 Manufacturing Census of Canada.³ Firm-level data on steam power adoption in the nineteenth century are available for other countries, notably including the Atack and Bateman (1999) samples from the nineteenth century U.S. Manufacturing Censuses. However given the objectives of this paper, the universal coverage in the CANIND71 database renders these data superior to the Atack and Bateman (1999) samples on account of two factors. Firstly, the empirical methodologies that are developed in Section 2.4 require estimation from various sub-samples, such as the sub-sample of firms using steam power. The large sub-samples that are available through the CANIND71 database ensure that the asymptotic properties of the estimators that are used can be reliably invoked. Secondly, as the main objective of this paper is to measure the aggregate welfare effect of steam power adoption, the universal coverage in the CANIND71 database minimizes the vulnerability of the results to sampling variation in estimation.

For the purposes of my analysis I exclude a number of industries that were included

³The history of the CANIND71 database is documented in Bloomfield and Bloomfield (1989). While the CANIND71 database is comprehensive, there are measurement issues in the raw manuscript data that must be reconciled prior to analysis. In particular, Inwood (1995) notes that multi-process establishments were occasionally decomposed by enumerators into multiple manuscript entries based on the distinct industrial activities carried out by the firm. Inwood (1995) finds that approximately 10 percent of the census manuscript entries were establishments that had been decomposed by enumerators. Using the data available in the CANIND71 database, I follow Inwood's (1995) procedure for identifying and reconstituting multi-process establishments. Inwood (1995, p. 364) uses the following reconstitution strategy: "The criteria for combining two or more entries into a larger and more complex firm are that they share a proprietor name within an enumeration division, that they appear immediately adjacent to each other in the manuscript schedule, and that an examination of the personal schedule 1 for the immediate area does not reveal the presence of two potential proprietors with the same name." I follow the same strategy as Inwood (1995), except that I omit the final criterion as I do not have access to the micro-data from schedule 1 of the 1871 Census of Canada. Reconstitution of the CANIND71 database reduces the number of establishments from 45,070 to 43,098.

in the 1871 Census of Manufacturing, but would not be categorized as manufacturing industries under a contemporary classification. The excluded industries include the following SIC major industry classes: agricultural services, forestry, mining, construction, gas and water utilities, personal and business services, and trade (including repair services). Excluding these industries reduces the sample size to 36,902. A further reduction in the sample occurs from excluding 62 firms that were located in remote census districts.⁴ Finally, I exclude firms that did not report a positive value of value added, labor, or fixed capital, which results in a cleaned sample of 32,581 observations.

Table 2.1 summarizes firms' production characteristics across the different types of power used in Canadian manufacturing in 1871. The observation counts highlight the fact that only 8 percent of Canadian manufacturing establishments had adopted steam power in 1871.⁵ By comparison, Atack et al. (2008) estimate that 20.1 percent of U.S. manufacturing establishments used steam power in 1870. However, caution is warranted in this comparison as differences in the enumeration practices between the Canadian and the U.S. censuses overstate the magnitude of this difference. In particular, the U.S. census did not enumerate establishments with revenue less than \$500, whereas Canadian enumerators were instructed to enumerate all establishments regardless of size, (Inwood and Keay, 2008). When the Canadian sample is censored to exclude firms falling below the U.S. census revenue threshold, the fraction of firms using steam power increases to 11 percent.⁶

⁴Remote census districts were identified as those having fewer than ten manufacturing establishments. These census districts were all located in remote northern regions of Ontario and Quebec.

⁵The classification of firms using steam power include 137 firms that reported using both steam and water power. All other firms in the sample were recorded as using only one of the five different types of power.

⁶For the purposes of converting the U.S. revenue threshold to Canadian dollars I use the official currency conversion for 1870, with 1 Canadian dollar being equivalent to 1.1492 U.S. dollars as reported in the online data set from Jacks and Pendakur (2010).

Table 2.1: Production Profile by Type of Power Used

	Hand	Water	Steam	Animal	Wind	Full Sample
Employees	3.20 (0.0594)	6.03 (0.271)	20.59 (0.933)	4.39 (0.186)	1.62 (0.154)	5.19 (0.104)
Fixed capital	711 (21.77)	3,736 (165.5)	11,534 (575.3)	1,154 (67.80)	591 (98.25)	2,189 (60.97)
Revenue	2,809 (70.79)	9,801 (419.6)	30,361 (1,773)	3,138 (200.0)	1,183 (243.5)	6,409 (177.4)
Value Added	1,360 (30.86)	3,098 (142.8)	14,279 (803.0)	1,840 (99.75)	399 (56.02)	2,755 (76.30)
Observations	22,617	6,379	2,630	905	50	32,581

Standard errors for the mean estimates are in parentheses. Fixed capital, revenue, and value added are in nominal 1871 Canadian dollars.

The most striking pattern in Table 2.1 is the positive correlation between firm size and the use of steam power. The average number of workers employed by a firm using steam power was roughly four times that of the overall average for the manufacturing sector. Furthermore, the value of fixed capital, revenue, and value added were approximately five times higher for firms with steam power. These findings are consistent with Atack et al. (2008), who also find a positive relationship between firm size and steam power adoption in the U.S. manufacturing sector during this era. The scale dominance of steam powered firms is illustrated dramatically in Figure 2.1. This figure plots the empirical cumulative distribution functions (CDFs) for the three most common types of power used by Canadian manufacturing firms in 1871: hand, water, and steam. The percentage of firms operating below any level of output is always highest for hand, followed by water, and then steam powered firms. This ordering of the three main technologies in the distribution of firm output is replicated in the theoretical model developed in Section 2.3.

The final set of empirical results in this section explores the relationship between

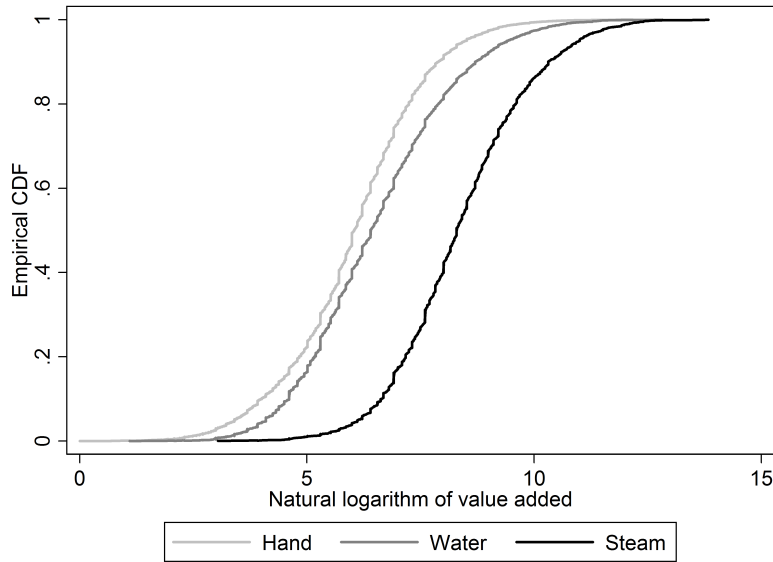


Figure 2.1: Empirical CDF of Natural Logarithm of Value Added by Type of Power Used

technological adoption and firm-level productivity. Table 2.2 presents results from several ordinary least squares (OLS) regressions of labor productivity on steam and water power adoption. The dependent variable is the natural logarithm of labor productivity (value added divided by employment weighted by months in operation). Value added is measured as gross sales minus the value of raw material inputs. The main coefficients of interest are associated with the indicator variables for firms' use of steam and water power. Column (1) reports estimates from a regression model including the main variables of interest and indicator variables that control for industry-specific effects. Column (2) adds controls for small firms (i.e. those having five or fewer employees), the gender composition of labor, the population in the firm's census district, and indicator variables for the firm's province of residence. Finally, column (3) includes an additional control variable for capital-intensity. Capital intensity is measured by the capital-labor ratio, and is calculated as the natural logarithm of fixed capital divided by employment weighted by months in operation.

Table 2.2: Regression Analysis of Labor Productivity for Canadian Manufacturing in 1871

Variable	(1)	(2)	(3)
Steam Power	0.636*** (0.0163)	0.489*** (0.0176)	0.118*** (0.0174)
Water Power	0.526*** (0.0134)	0.490*** (0.0136)	0.0256* (0.0147)
Five or Fewer Employees		-0.0939*** (0.0126)	-0.131*** (0.0119)
Male Employees/Employees		0.614*** (0.0201)	0.466*** (0.0188)
ln(District Population)		0.159*** (0.00832)	0.109*** (0.00780)
ln(Capital/Labor)			0.252*** (0.00421)
Province Controls	No	Yes	Yes
SIC Industry Controls	Yes	Yes	Yes
Observations	32,581	32,581	32,581
R-squared	0.158	0.206	0.306

Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Coefficient estimates for province and industry indicator variables are not reported but are available upon request.

The regression results in Table 2.2 indicate that steam and water power were correlated with higher labor productivity in the Canadian manufacturing sector in 1871, holding constant the other control variables in the regression analysis. The coefficient on the steam power variable is positive, and statistically significant at the 1 percent level in each specification. The coefficient on the water power variable is positive, and statistically significant at the 1 percent level in specifications (1) and (2), and positive and statistically significant at the 10 percent level in specification (3). The magnitude of water and steam power coefficients drops considerably in specification (3). This is not surprising as steam and water power were embodied in physical capital. The coefficients for these variables in column (3) therefore capture the correlation between these technologies and labor productivity as distinct from other forms of capital investment. Atack et al. (2008) note that relative to water power, steam was a flexible power source that had the clear advantage of not being tied to a geographic location. This logic provides a reasonable explanation for the finding that steam is more highly correlated with productivity, after controlling for capital intensity.

The coefficients for the other control variables in Table 2.2 are consistent with the literature on the North American manufacturing sector during this era.⁷ The negative and highly significant coefficient for the small firm indicator variable is similar to the results of Inwood and Keay (2012), who find evidence of returns to scale in Canadian manufacturing during this era. The positive and highly significant coefficient on the variable measuring the fraction of male employees is consistent with the findings of Goldin and Sokoloff (1982), who find that female wages were approximately 50 percent lower than male wages during the mid-nineteenth century in the U.S. manu-

⁷The coefficient estimates for the small firm indicator variable, the fraction of male employees, and the natural log of district population are statistically significant at the 1 percent level in specifications (2) and (3).

facturing sector. Finally, the positive and highly significant coefficient for the census district population variable is analogous to the results of Inwood and Keay (2005), who find evidence of a positive correlation between market size and productivity for the Canadian manufacturing sector during this era.

The empirical analysis in this section can be summarized by three main results. First, a relatively small fraction of firms used steam power in the Canadian manufacturing sector in 1871. Second, the firms using steam power were large, with value added that was approximately five times the average for the Canadian manufacturing sector. Third, steam and water power adoption were both positively correlated with labor productivity. For both technologies this positive correlation persists when capital intensity is controlled for, however the magnitude and statistical significance of the results are strongest for steam.

The analysis in this section does not identify the causal effect of steam power adoption on productivity, and the analysis is also not capable of quantifying the effect of steam power on welfare in the aggregate economy. In Sections 2.3 and 2.4 I develop and estimate a model of heterogeneous firms and technological adoption to achieve these objectives.

2.3 Theory

This section develops a model of technological adoption with monopolistically competitive firms that are heterogeneous with respect to their productivity. The theoretical framework extends the model of Bustos (2011) by allowing for multiple technologies and by deriving a welfare statistic that can be estimated using firm-level data. In the model, firms make a discrete choice to operate using one of several different technologies. These technologies vary from lowest to highest productivity (output

per worker). Despite the superiority of the most productive technology, firm-level heterogeneity and differences in the cost of adoption results in each of the production technologies being used by some fraction of firms in equilibrium. The ultimate objective of this section of the paper is to use the model to derive a statistic that measures the welfare gains from the introduction and adoption of the steam power. In Section 2.4, I illustrate how this welfare statistic can be estimated using firm-level data. For concreteness, the model below refers to firms' adoption of steam power, but the framework can be applied to estimate the gains from the adoption of other technologies.

2.3.1 Model

I consider an economy with a unit measure of identical households with CES utility:

$$U = \left(\int_{\omega \in \Omega} q(\omega)^{\frac{\sigma-1}{\sigma}} d\omega \right)^{\frac{\sigma}{\sigma-1}}, \quad \sigma > 1, \quad \rho \equiv \frac{\sigma-1}{\sigma}, \quad (2.1)$$

where $q(\omega)$ is the quantity of variety ω demanded by the representative household, σ is the elasticity of substitution between varieties, and Ω is the endogenous set of varieties produced. Each household has one unit of labor that is supplied inelastically to firms. Labor is defined as the numeraire, and the wage is normalized to one. Households also hold an equal share in the profits of producers.

Firms use the following technologies in production:

$$q(\varphi, \gamma) = \gamma \varphi l, \quad \gamma \in \{\gamma_1, \gamma_2, \gamma_3\}, \quad \gamma_3 > \gamma_2 > \gamma_1 = 1 \quad (2.2)$$

The parameter γ differentiates the production technology across the three different power sources. That is, the firms' choice to use hand, water, or steam power is for-

malized by the associated choice of production technology $q(\varphi, 1)$, $q(\varphi, \gamma_2)$, or $q(\varphi, \gamma_3)$ respectively. Steam power's superiority over water and hand power is captured in the assumption that $\gamma_3 > \gamma_2 > 1$, where I have normalized the productivity parameter for hand power to one for notational convenience.

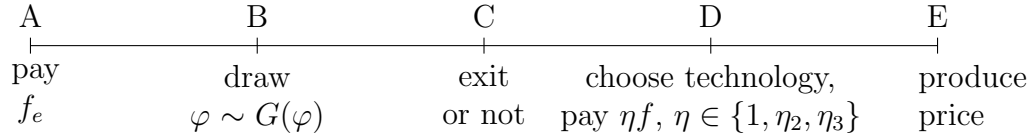


Figure 2.2: Firms' Timeline

The production functions in equation (2.2) also depend on the parameter φ , which is the firm's idiosyncratic productivity type. The firm's productivity type is realized through the process summarized in Figure 2.2. At stage A, a prospective firm chooses to enter the industry by paying a fixed cost f_e . All fixed costs are denoted in units of labor. After paying the entry cost, firms draw their productivity type, φ , from a Pareto CDF, $G(\varphi) = 1 - \varphi^{-\theta}$. I assume that the shape parameter of the Pareto distribution, θ , satisfies $\theta > \sigma - 1$, which ensures that the ex-post average productivity of firms is finite. Upon realizing their type, firms face a decision at stage C to exit or remain in the industry. At stage D, firms choose a power type and pay a technology specific fixed cost of production. The firms' decision to use hand, water, or steam power is associated with a fixed cost of production equal to f , $\eta_2 f$, or $\eta_3 f$ respectively. I assume that $\eta_3 > \eta_2 > 1$ which implies that the cost of using steam is greater than for water, and that the cost of using water is greater than for hand power. Finally, at stage E firms choose their price and production level to maximize profits.

Solving the representative household's utility maximization problem yields the

following CES demand function for variety ω :

$$q(\omega) = P^{\sigma-1}p(\omega)^{-\sigma}, \quad P \equiv \left(\int_{\omega \in \Omega} p(\omega)^{1-\sigma} d\omega \right)^{\frac{1}{1-\sigma}} \quad (2.3)$$

where P is the CES price index.

Next I consider a firm's problem by working backward from stage E. Monopolistic competition and CES demand implies that the price chosen by a firm of type φ is: $p(\varphi, \gamma) = 1/(\rho\gamma\varphi)$. Combining the firm's pricing rule with household demand yields the following revenue and profit function:

$$r(\varphi, \gamma) = P^{\sigma-1}p(\varphi, \gamma)^{1-\sigma}, \quad \pi(\varphi, \gamma) = \frac{r(\varphi, \gamma)}{\sigma} - \eta f \quad (2.4)$$

At stage D, firms choose a power type which affects profitability through the productivity parameter, γ , and through the fixed cost, $\eta \in \{1, \eta_2, \eta_3\}$. To match the empirical regularities presented in Section 2.2, I consider an equilibrium where all three technologies are used by some fraction of firms, and where steam is only used by the largest and most productive firms. Define φ_3 as the cut-off, such that a firm with this productivity level is just indifferent with respect to using steam or water power. That is, $\pi(\varphi_3, \gamma_3) = \pi(\varphi_3, \gamma_2)$, which from the profit function in equation (2.4) implies:

$$\varphi_3 = \frac{(\sigma f)^{\frac{1}{\sigma-1}}}{\rho P} \left(\frac{\eta_3 - \eta_2}{\gamma_3^{\sigma-1} - \gamma_2^{\sigma-1}} \right)^{\frac{1}{\sigma-1}} \quad (2.5)$$

Next consider the subset of firms choosing water power at stage D. Define φ_2 as the cut-off, such that a firm with this productivity level is just indifferent with respect to using water or hand power. That is, $\pi(\varphi_2, \gamma_2) = \pi(\varphi_2, 1)$, which from the profit

function in equation (2.4) implies:

$$\varphi_2 = \frac{(\sigma f)^{\frac{1}{\sigma-1}}}{\rho P} \left(\frac{\eta_2 - 1}{\gamma_2^{\sigma-1} - 1} \right)^{\frac{1}{\sigma-1}} \quad (2.6)$$

Active firms that are not productive enough to use water or steam choose to operate with hand power. Define φ_1 as the operating cut-off, such that a firm with this productivity level is just indifferent between operating with hand power and exiting after learning its productivity type. By definition a firm that draws φ_1 makes zero profit, $\pi(\varphi_1, 1) = 0$. From the profit function in equation (2.4) it follows that $\varphi_1 = (\sigma f)^{\frac{1}{\sigma-1}} / (\rho P)$. Combining this expression with equations (2.5) and (2.6) yields the following relative cut-off equations:⁸

$$\Delta_{31} \equiv \frac{\varphi_3}{\varphi_1} = \left(\frac{\eta_3 - \eta_2}{\gamma_3^{\sigma-1} - \gamma_2^{\sigma-1}} \right)^{\frac{1}{\sigma-1}}, \quad \Delta_{21} \equiv \frac{\varphi_2}{\varphi_1} = \left(\frac{\eta_2 - 1}{\gamma_2^{\sigma-1} - 1} \right)^{\frac{1}{\sigma-1}} \quad (RC)$$

Define $\tilde{\pi}$ as the expected profit of an active firm, $\tilde{\pi} = E(\pi|\varphi \geq \varphi_1)$. Given the profit function in equation (2.4), and the relative cut-off equations (RC), the expected profit of an active firm can be derived and is given by the following *zero-cut-off-profit* condition:

$$\tilde{\pi} = \frac{f(\sigma - 1)}{\theta - \sigma + 1} (1 + (\eta_2 - 1)\Delta_{21}^{-\theta} + (\eta_3 - \eta_2)\Delta_{31}^{-\theta}) \quad (ZCP)$$

Finally at stage A firms will enter the industry up until the point that expected profits are equal to the fixed cost of entry, $(1 - G(\varphi_1))\tilde{\pi} = f_e$. This free entry

⁸In an equilibrium where all three technologies are used, the most productive firms use steam, and the least productive firms use hand power; it must be the case that $\varphi_1 < \varphi_2 < \varphi_3$. From equation (RC), it can be seen that this ordering of the cut-offs requires assuming that $\eta_2 - 1 > \gamma_2^{\sigma-1} - 1$ and $(\gamma_2^{\sigma-1} - 1)(\eta_3 - \eta_2) > (\gamma_3^{\sigma-1} - \gamma_2^{\sigma-1})(\eta_2 - 1)$. I impose these assumptions to derive an equilibrium that is consistent with the empirical regularities presented in Section 2.2.

condition can be rewritten:

$$\tilde{\pi} = fe / (1 - G(\varphi_1)) = f_e \varphi_1^\theta \quad (FE)$$

The equilibrium operating cut-off, φ_1 , can be solved by combining equations (ZCP) and (FE):

$$\varphi_1 = \left(\frac{f(\sigma - 1)}{fe(\theta - \sigma + 1)} [1 + (\eta_2 - 1)\Delta_{21}^{-\theta} + (\eta_3 - \eta_2)\Delta_{31}^{-\theta}] \right)^{\frac{1}{\theta}} \quad (2.7)$$

Define φ'_1 as the equilibrium operating cut-off when steam power is not available in the economy. Solving the model without steam power yields:

$$\varphi'_1 = \left(\frac{f(\sigma - 1)}{fe(\theta - \sigma + 1)} [1 + (\eta_2 - 1)\Delta_{21}^{-\theta}] \right)^{\frac{1}{\theta}} = \lim_{\eta_3 \rightarrow \infty} \varphi_1 = \lim_{\gamma_3 \rightarrow \gamma_2} \varphi_1 \quad (2.8)$$

Intuitively, when the cost of steam power adoption becomes prohibitively high, or when the productivity increase associated with steam power adoption is small relative to water power, the equilibrium operating cut-off approaches its value in the economy without steam power.

2.3.2 Measuring the Welfare Gains from Technological Adoption

In this section I derive a statistic that summarizes how the introduction and adoption of steam power affects welfare.⁹ The primary objective of this exercise is to derive a representation of the welfare statistic that can be estimated from firm-level data.

⁹The methodology developed in this section draws from the literature on estimating the welfare gains from international trade, including Melitz and Redding (2014), Arkolakis et al. (2012), Feenstra (2010).

To derive this statistic, I compare the equilibrium of the model with hand, water, and steam power, to an equilibrium where only hand and water power are available to firms. Under CES preferences, welfare is measured by the real wage, which is equal to the reciprocal of the price index when the nominal wage is normalized to one, $\mathbb{W} = w/P = 1/P$. I denote welfare and other values in the equilibrium without steam with primes, $\mathbb{W}' = 1/P'$.

Proposition 1: The change in welfare from the introduction and adoption of steam power can be measured by the increase in the equilibrium operating cut-off, φ_1 . That is:

$$\mathbb{W}/\mathbb{W}' = P'/P = \varphi_1/\varphi_1' = \left(\frac{1 + (\eta_2 - 1)\Delta_{21}^{-\theta} + (\eta_3 - \eta_2)\Delta_{31}^{-\theta}}{1 + (\eta_2 - 1)\Delta_{21}^{-\theta}} \right)^{\frac{1}{\theta}} \quad (2.9)$$

Proof: see Appendix A.

Estimating the effect of steam power on welfare requires an estimate of the shape parameter of the Pareto distribution, θ , and an estimate of the term in large parenthesis in the rightmost expression in equation (2.9). With the objective of finding an expression for this term that can be estimated from firm-level data, I use the labor market clearing condition to solve for the equilibrium measure of firms, M , that are active in the economy with steam power:

$$M = \frac{\theta - \sigma + 1}{\theta \sigma f} \left(1 + (\eta_2 - 1)\Delta_{21}^{-\theta} + (\eta_3 - \eta_2)\Delta_{31}^{-\theta} \right)^{-1}$$

$$\iff 1 + (\eta_2 - 1)\Delta_{21}^{-\theta} + (\eta_3 - \eta_2)\Delta_{31}^{-\theta} = \frac{\theta - \sigma + 1}{\theta \sigma f M} \quad (2.10)$$

Equation (2.10) is recognizable as the numerator of term in the parenthesis in equation (2.9). Substituting this expression and imposing goods market clearing

$R = wL = 1$, where R is aggregate revenue, yields:

$$\mathbb{W}/\mathbb{W}' = \left\{ R / \left(\frac{\theta \sigma f M}{\theta - \sigma + 1} [1 + (\eta_2 - 1) \Delta_{21}^{-\theta}] \right) \right\}^{\frac{1}{\theta}} \quad (2.11)$$

The final step in deriving a welfare statistic that can be estimated from firm-level data is to find a convenient expression for the term $(\theta \sigma f M)(1 + (\eta_2 - 1) \Delta_{21}^{-\theta}) / (\theta - \sigma + 1)$ in equation (2.11). With this objective in mind, define R_1 and R_2 as the aggregate revenue produced by firms using hand and water power respectively in the economy with steam power. Furthermore, define R_{-3} as the aggregate revenue that would have been realized by steam powered firms, had they been forced to operate with the second most productive technology (i.e. water power), holding constant the equilibrium productivity cut-offs, price index, and measure of firms at their values in the economy with steam power. Given these definitions, we have:

$$\begin{aligned} & R_1 + R_2 + R_{-3} \\ &= \int_{\varphi_1}^{\varphi_2} r(\varphi, 1) \frac{g(\varphi)}{1 - G(\varphi_1)} M d\varphi + \int_{\varphi_2}^{\varphi_3} r(\varphi, \gamma_2) \frac{g(\varphi)}{1 - G(\varphi_1)} M d\varphi + \int_{\varphi_3}^{\infty} r(\varphi, \gamma_2) \frac{g(\varphi)}{1 - G(\varphi_1)} M d\varphi \\ &= \frac{\theta \sigma f M}{\theta - \sigma + 1} (1 - \Delta_{21}^{-\theta + \sigma - 1}) + \frac{\theta \sigma f M}{\theta - \sigma + 1} (\Delta_{21}^{-\theta + \sigma - 1} - \Delta_{31}^{-\theta + \sigma - 1}) \gamma_2^{\sigma - 1} + \frac{\theta \sigma f M}{\theta - \sigma + 1} \Delta_{31}^{-\theta + \sigma - 1} \gamma_2^{\sigma - 1} \\ &= \frac{\theta \sigma f M}{\theta - \sigma + 1} (1 + (\eta_2 - 1) \Delta_{21}^{-\theta}) \end{aligned}$$

Note that even though the expression above integrates over the range of firms using steam power, $\varphi \geq \varphi_3$, the definition of R_{-3} imposes that the revenue of these firms be calculated in a counterfactual environment in which they produce with water power, $\gamma = \gamma_2$. Combining this expression with equation (2.11) yields the following welfare equation.

$$\mathbb{W}/\mathbb{W}' = (R / (R_1 + R_2 + R_{-3}))^{\frac{1}{\theta}}, \quad (2.12)$$

In Section 2.4, I illustrate how this welfare equation can be estimated using firm-level data.

2.4 Empirical Application

2.4.1 Econometric Model

In this section I develop an econometric model for estimating the welfare statistic given by equation (2.12). The quantities R , R_1 , and R_2 , can be estimated by aggregate value added, and the aggregate value added produced by firms using hand and water power respectively. The empirical analogue to the theoretical quantity defined by R_{-3} is as follows:

$$\hat{R}_{-3} = \sum_{i=1}^{N_3} \hat{r}_i(\varphi_i, \gamma_2), \quad (2.13)$$

where N_3 is the number of firms using steam power, and $\hat{r}_i(\varphi_i, \gamma_2)$ is the revenue that would have been produced by a steam powered firm, had they been forced to use water power.

An empirical methodology for estimating $\hat{r}_i(\varphi_i, \gamma_2)$ can be derived from the theoretical model. Taking the natural logarithm of the revenue equation (2.4) after substituting in the monopolistically competitive price, and considering only the subset of firms using steam or water power yields:

$$\ln(r(\varphi, \gamma)) = (\sigma - 1)\ln(\rho P \gamma_2) + 1_{\gamma_3}(\sigma - 1)\ln\left(\frac{\gamma_3}{\gamma_2}\right) + (\sigma - 1)\ln(\varphi), \quad (2.14)$$

where 1_{γ_3} is an indicator function that takes a value of one if the firm uses steam and

zero otherwise. The empirical specification corresponding to equation (2.14) is:

$$y_i = \beta_0 + \beta_1 1(\text{steam}_i) + u_i, \quad (2.15)$$

where y_i is the natural logarithm of value added, $1(\text{steam}_i)$ is a steam power indicator variable that takes a value of one if firm i uses steam and zero otherwise, and regression equation (2.15) is defined over the sub-sample of firms using steam or water power. The regression coefficients and error term are defined as follows: $\beta_0 = (\sigma - 1)\ln(\rho P \gamma_2)$, $\beta_1 = (\sigma - 1)\ln(\gamma_3/\gamma_2)$, and $u_i = (\sigma - 1)\ln(\varphi_i) + \epsilon_i$. The error term, u_i , includes firm i 's unobserved productivity type, φ_i , and the random variable ϵ_i which is assumed to be i.i.d. across firms. If a firm's productivity type is positively correlated with the adoption decision, then the coefficient β_1 will be biased upwards if estimated by OLS regression. A consistent estimate of β_1 can be obtained using instrumental variables (IV) estimation. In Section 2.4.2 I discuss the instruments I use to estimate equation (2.15) by IV estimation.

Once β_1 has been estimated, $\hat{r}_i(\varphi_i, \gamma_2)$ can be calculated for the subset of firms using steam power as follows:

$$\hat{r}_i(\varphi_i, \gamma_2) = \exp(y_i - \hat{\beta}_1 1(\text{steam}_i)), \quad (2.16)$$

where $\hat{\beta}_1$ is the estimated value of β_1 and the remaining terms in this expression are defined in equation (2.15). Given an estimate of $\hat{r}_i(\varphi_i, \gamma_2)$, equation (2.13) can in turn be used to calculate \hat{R}_{-3} .

The final parameter required to estimate the welfare equation given by (2.12) is an estimate of the shape parameter of the Pareto distribution, θ . I now show how an estimate of θ can be obtained by estimating the parameters of the distribution

of value added for the sub-sample of firms that had adopted steam power in 1871. Define $F(r)$ as the CDF of value added for the subset of firms having adopted steam power. In Appendix A I show that $F(r)$ follows a Pareto distribution with shape parameter $\theta/(\sigma - 1)$. Therefore, for the subset of firms using steam power, the CDF $F(r)$ can be written:

$$\ln(r) = \ln(r(\varphi_3, \gamma_3)) - \frac{\sigma - 1}{\theta} \ln(1 - F(r)), \quad (2.17)$$

where $r(\varphi_3, \gamma_3)$ is the revenue of a steam powered firm with productivity equal to the steam cut-off, φ_3 . The empirical specification corresponding to equation (2.17) is:

$$y_i = \zeta_0 + \zeta_1 \ln(1 - \hat{F}_i) + \epsilon_i \quad (2.18)$$

where y_i is the natural logarithm of value added, \hat{F}_i is the empirical CDF, and regression equation (2.18) is defined over the subset of firms using steam power. I follow Head et al. (2014) in estimating equation (2.18) by the QQ regression methodology introduced by Kratz and Resnick (1996). Defining $i = 1$ as the firm with the minimum value added and $i = n$ as the firm with the maximum value added in the sub-sample, the empirical CDF is defined as $\hat{F}_i = (i - 0.3)/(n + 0.4)$. The coefficient of interest is $\zeta_1 = -(\sigma - 1)/\theta$. Identifying θ requires an estimate of σ , the elasticity of substitution between varieties. Hsieh and Klenow (2009) note that estimates of σ in the empirical trade and industrial organization literature typically range from 3 to 10 for the manufacturing sector. Ziebarth (2013) assumes that $\sigma = 3$ in his analysis of misallocation and productivity of the U.S. manufacturing sector in the nineteenth century. For my analysis, I follow Melitz and Redding (2013) and Head et al. (2014) in using a benchmark parameterization of $\sigma = 4$ and explore the robustness of this

assumption in the sensitivity analysis.

Once the parameter θ and the expressions R , R_1 , R_2 , and R_{-3} have been estimated, the welfare effect of the introduction and adoption of steam power can be calculated directly using the statistic defined by equation (2.12).

2.4.2 Identification

In theory it is straightforward to apply the estimation strategy that is presented in Section 2.4.1. However in practice, finding an exogenous source of variation that can be used to derive a consistent estimate of β_1 is challenging. In the case of steam power, it is possible to use exogenous variation in geography to identify the parameter β_1 . In particular I use average agricultural soil quality and precipitation at the district-level as instrumental variables to estimate the effect of steam power adoption on firm-level productivity. This identification strategy draws on the theory proposed by Rothbarth (1946), and Habakkuk (1962), that agricultural land quality can explain regional differences in the pace of industrialization. In particular, these authors argue that the superior land endowment in the U.S. raised the opportunity cost of labor and encouraged the substitution of capital for labor to a greater extent in the U.S. than in the UK during the nineteenth century. In the Canadian context, McCallum (1980) has also argued that regional differences in industrial growth can be explained by agricultural land quality, although his argument relies on an alternative explanation. McCallum (1980) argues that regions of superior land quality were quicker to industrialize as a result of the capital formation financed by regional wealth accumulated in the agricultural sector during the nineteenth century. Based on the arguments of these authors, it is reasonable to hypothesize a positive correlation between firm-level steam power adoption and exogenous variation in average

agricultural soil quality and precipitation levels, insofar as these measures proxy for regional agricultural endowments.

To construct a measure of average soil quality at the district-level I use geographic information systems (GIS) analysis and digital data from the Canada Land Inventory (CLI).¹⁰ The CLI digital maps classify soil capability for agriculture using a seven-point descending scale. Class 1 is the highest agricultural soil quality classification and is described as having “no significant limitations in use for crops”, (Agriculture and Agri-Food Canada, 2015). Class 7 is the lowest quality soil classification and is deemed to “have no capacity for arable culture or permanent pasture”, (Agriculture and Agri-Food Canada, 2015). Using GIS analysis, I estimate average agricultural soil quality at the 1871 Canadian census district-level.¹¹ I also use GIS analysis to estimate the average annual precipitation within the census districts as a second instrument to be used in IV estimation. It is important to include a measure of average annual precipitation as agricultural production depends both on soil quality and precipitation. Geo-spatial precipitation data is sourced from WorldClim Global Climate Data, which is publicly available on the internet.¹² The WorldClim data

¹⁰The CLI was an initiative administered by the Canadian federal government’s Agricultural Rehabilitation and Development Act of 1961. The ninety year lag between period of the analysis and the CLI initiative limits the accuracy of CLI data in approximating soil quality in the late nineteenth century. Despite this limitation, the CLI data provide a reasonable proxy for soil quality in 1871 as the classification was based on the potential of soils for agricultural production, rather than its observed agricultural productivity in 1961.

¹¹Josh MacFadyen (University of Saskatchewan) provided the geo-spatial data for the 1871 Canadian census districts for the provinces of Ontario and Quebec. For the provinces of New Brunswick and Nova Scotia, I use the 1911 Canadian census districts, geo-edited to match the 1871 census district boundaries. The 1911 Canadian census districts are sourced from the Canadian Century Research Infrastructure project at the University of Ottawa, Canada. When calculating averages at the census-district level, I impose a two kilometer buffer around the census boundaries as a robustness measure against imprecision in the geo-referencing for the 1871 census district spatial data. While most of the census districts represented relatively large geographic areas, a number of Canadian cities were defined as self-contained census districts and consequently represented a much smaller geographic area. To address this issue, I merge cities that were self-contained districts with a surrounding district to ensure greater consistency in the geographic definition of the districts.

¹²The time lag between the period of the analysis and the precipitation data limits the accuracy of WorldClim estimates in approximating precipitation patterns in the late nineteenth century. How-

provides high geo-spatial resolution estimates of average annual precipitation over the period 1950-2000.

The instrumental variables strategy proposed above will produce a consistent estimate of β_1 if two criteria are satisfied. The first criteria requires that the instruments are highly correlated with steam power adoption. In theory, the arguments of Rothbarth (1946), Habakkuk (1962), and McCallum (1980) suggest that the instruments should be highly correlated with adoption. Empirically, this first criteria is tested by the first-stage F-statistic in two-stage least squares, which measures the correlation between the instruments and steam power adoption at the firm-level. The second criteria requires that the instruments satisfy the ‘exclusion restrictions’. That is, the instruments must not be correlated with any components in the error term, u_i , that are correlated with the firm-level value added.

There are several potential violations of the exclusion restrictions that must be considered. To address these challenges I augment regression specification (2.15) with several control variables as follows:

$$y_i = \beta_0 + \beta_1 1(\text{steam}_i) + \mathbf{X}_i \beta + u_i, \quad (2.19)$$

where \mathbf{X}_i is a vector of control variables. These control variables are included to address several possible challenges to the exclusion restrictions. For example, districts having the richest agricultural endowments were likely more densely populated in 1871. This higher population density may have increased firm-level output through agglomeration effects in local labor markets. To address this issue the vector of control variables includes the population density of each district as sourced from the 1871

ever, given that the WorldClim data were collected over a fifty-year period in the twentieth century, the data can be expected to reflect long run trends and provide a reasonable proxy for precipitation in 1871. The WorldClim data are publically available at the following url: <http://www.worldclim.org/>

Census of Canada. Districts with the best agricultural soils in Canada also likely had greater access to railroads in the late nineteenth century. In theory, rail access may have increased firm-level output in manufacturing through market access and external economies of scale. This issue is addressed by including a rail access indicator variable that takes a value of one if the district was serviced by a railroad. The data for the rail access variable is sourced from the 1863 Canadian railroad maps produced by Andreae (1996). Districts with high precipitation were likely to have had a higher concentration of rivers. Access to rivers may be correlated with firm-level output as many industrial activities benefited from access to running water. The vector of controls includes a variable that measures the natural logarithm of average stream flow accumulation in each district. I estimate average stream flow accumulation by GIS using a digital elevation model (DEM) with data sourced from the NASA Shuttle Radar Topography Mission (SRTM).¹³ There may also have been direct benefits to manufacturing firms located in districts with rich agricultural endowments. This may have been especially true for flour and grist mills which represented an important industry in the Canadian economy during this era. This issue is partially addressed by using a value added measure of output, which nets out the value of agricultural and other intermediate inputs. However, there may have been additional benefits for firms that located in close proximity to high quality agricultural inputs that are not fully captured in the value of intermediate inputs. I address this issue by including a control variable measuring the total wheat produced per acre in each district.¹⁴ Finally, the vector of controls also includes indicator variables for each SIC major industry group, and provincial indicator variables. These indicator variables control for industry specific technological differences and inter-provincial variation in input

¹³SRTM DEM data was downloaded from the Consultative Group for International Agricultural Research - Consortium for Spatial Information (CGIAR-CSI) website: <http://www.cgiar-csi.org/>

¹⁴These data are sourced from the 1871 Census of Canada.

and output prices.

2.4.3 Estimation Results

The analysis in this section uses firm-level data from the 1871 Canadian Census of Manufacturing sourced from the CANIND71 database that was described in Section 2.2. These data are supplemented with the variables described in Section 2.4.2. The first step required to estimate the welfare statistic involves estimating equation (2.19) using the sub-sample of firms operating with either water or steam power. I estimate equation (2.19) by OLS, and by two variants of instrumental variables estimation.

The first IV estimator used is the standard two-stage least squares methodology (henceforth, 2SLS). The second IV estimator is the *Procedure 18.1* approach proposed by Wooldridge (2002). The first step in Wooldridge’s *Procedure 18.1* estimator (henceforth, IV-WP) involves regressing the steam power indicator variable on the full set of instruments and the control variables from the equation (2.19) using a probit regression.¹⁵ The fitted probabilities are then obtained from the probit regression results and are used as an instrumental variable in estimating equation (2.19) by two-stage least squares. The IV-WP estimator is more efficient than standard two-stage least squares estimation, however this result relies on the probit regression model being correctly specified in the first step, (Wooldridge, 2002).¹⁶ Importantly, the IV-WP estimator remains consistent and the standard errors of the coefficient estimates are asymptotically valid, even when probit regression is misspecified.

Table 2.3 reports the OLS, 2SLS, and IV-WP regression results along with the

¹⁵More generally, the first step in Wooldridge’s *Procedure 18.1* can be estimated by any binary response model using maximum likelihood.

¹⁶In the context of the theoretical model developed in this paper, a probit specification of the binary response model would be correctly specified if I assumed that ex-ante firm-level productivity followed the log-normal distribution instead of the Pareto distribution. Head et al. (2014) argue that there are theoretical and empirical justifications for using the log-normal distribution as an alternative to the Pareto distribution when modeling firm-level productivity.

Table 2.3: Regression of the Natural Logarithm of Value Added on Steam Power Adoption

Variable/Estimate	OLS	2SLS	IV-WP
Steam Power	1.123*** (0.0408)	0.600*** (0.168)	0.423*** (0.142)
Population Density	1.636*** (0.0757)	1.785*** (0.0897)	1.836*** (0.0856)
Rail Access	0.0638** (0.0324)	0.0896*** (0.0332)	0.0983*** (0.0332)
ln(Stream Flow)	0.172*** (0.0156)	0.176*** (0.0158)	0.178*** (0.0159)
Wheat/Acre	0.271*** (0.0493)	0.306*** (0.0497)	0.318*** (0.0495)
γ_3/γ_2	1.454*** (0.0198)	1.221*** (0.0686)	1.151*** (0.0545)
$R/(R_1 + R_2 + R_{-3})$	1.393	1.233	1.168
Observations	9,009	9,009	9,009
R-squared	0.418	0.406	0.397
Kleibergen-Paap rk Wald F-statistic		231.4	

*** p<0.01, ** p<0.05, * p<0.1. For γ_3/γ_2 the asterisks represent statistical significance for a test of the null hypothesis that this ratio is equal to one. Robust standard errors in parentheses. Indicator variables for SIC major industry groups and province of establishment are included in all specifications. Coefficient estimates for these variables are not reported but are available upon request. The estimate of γ_3/γ_2 is calculated using a parameterized value of $\sigma = 4$.

estimates of γ_3/γ_2 and the ratio $R/(R_1 + R_2 + R_{-3})$. The estimates of γ_3/γ_2 use the parameterized value of $\sigma = 4$, and are derived from the coefficient estimate for the steam power indicator variable, $\beta_1 = (\sigma - 1)\ln(\gamma_3/\gamma_2)$. The standard errors for the ratio γ_3/γ_2 are calculated using the delta method. To estimate the ratio $R/(R_1 + R_2 + R_{-3})$, the terms R , R_1 , and R_2 are calculated from the census data on aggregate value added, and the value added produced by firms using hand and water power respectively. The term R_{-3} is calculated from the census data on steam powered firms combined with the regression results, and using equations (2.13) and (2.16).

The results in Table 2.3 suggest a positive and statistically significant relationship between steam power adoption and value added. The coefficient estimates for the steam power indicator variable are positive and statistically significant at the 1 percent level in every column of Table 2.3. This is not surprising given the results that were reported in Section 2.2. More interesting is the decline in the estimates of β_1 and γ_3/γ_2 between OLS and the IV estimators. Recall that the ratio γ_3/γ_2 measures the gross increase in firm-level productivity from using steam power relative to water power. The 2SLS and IV-WP estimates of this parameter are 23 and 30 percentage points lower than the OLS estimate respectively. This is consistent with the notion that positive selection results in an upward bias in the OLS estimate of the effect of steam power on productivity. The causal interpretation of the IV estimates are that the gain from switching from water to steam power was a 22.1 and 15.1 percent increase in firm-level labor productivity for the 2SLS and IV-WP estimates respectively.¹⁷ While the magnitude of these estimates are large, they are not unreasonable when they are interpreted in light of the fact that the average value added of firms using steam power was more than four times that of water powered firms.

¹⁷These estimates are statistically significant at the 1 percent level.

The estimates for the ratio $R/(R_1 + R_2 + R_{-3})$ are also reported in Table 2.3. The 2SLS and IV-WP derived estimates of this ratio show a marked decline from the OLS estimates. I choose to use the IV-WP estimate of $R/(R_1 + R_2 + R_{-3})$ in calculating the welfare statistic as it is the most conservative of the estimates.

Table 2.4: Regression of Steam Power Adoption on Instruments and Control Variables

Variables	OLS	Probit
Agricultural Soil Quality	-0.0810*** (0.00438)	-0.306*** (0.0186)
Precipitation	0.000792*** (6.13e-05)	0.00328*** (0.000309)
Population Density	0.193*** (0.0241)	0.747*** (0.0994)
Rail Access	0.0400*** (0.00863)	0.230*** (0.0395)
ln(Stream Flow)	0.0151*** (0.00415)	0.0641*** (0.0161)
Wheat/Acre	-0.0633*** (0.0185)	-0.264*** (0.0615)
Province Controls	Yes	Yes
SIC Industry Controls	Yes	Yes
Observations	9,009	9,009
R-squared	0.361	

Robust standard errors in parentheses. Coefficient estimates for province and industry indicator variables are not reported but are available upon request.

For the 2SLS estimates, the Kleibergen-Paap rk Wald F-statistic is reported at the bottom of Table 2.3. The null hypothesis of weak-identification is soundly rejected at any reasonable level of significance. In Table 2.4 I also report the first-stage regression results for the 2SLS estimator, and the probit regression used in the IV-WP estimator. Both soil quality and precipitation are highly significant and of the expected sign, providing further evidence in support of the strength of the identification strategy. In particular, the results suggest a strong positive correlation between firm-level adoption

of steam power and regional agricultural endowments, as measured by soil quality and precipitation.

Table 2.5: QQ Regression Results and Estimates of the Welfare Effect of Steam Power Adoption

Estimate	All	Top 50	Top 25	Top 20	Top 15	Top 10	Top 5
$-(\sigma - 1)/\theta$	-1.362 (0.0110)	-0.977 (0.00546)	-0.811 (0.00607)	-0.766 (0.00629)	-0.717 (0.00627)	-0.658 (0.00641)	-0.574 (0.00803)
θ	2.202 (0.0177)	3.069 (0.0171)	3.701 (0.0277)	3.914 (0.0321)	4.184 (0.0366)	4.562 (0.0445)	5.229 (0.0732)
\mathbb{W}/\mathbb{W}'		1.052	1.043	1.041	1.038	1.035	1.030
Observations	2,630	1,349	658	526	408	282	133
R-squared	0.855	0.960	0.964	0.966	0.970	0.974	0.975

Standard errors are in parentheses. The estimates of the shape parameter, θ , and the welfare statistic, \mathbb{W}/\mathbb{W}' , are calculated using $\sigma = 4$, and the IV-WP derived estimate of $R/(R_1 + R_2 + R_{-3})$ from Table 2.3.

The next step required to estimate the welfare statistic is to estimate the shape parameter of the Pareto distribution, θ . The QQ regression estimation results of equation (2.18) are reported in Table 2.5. I follow Head et al. (2014) in reporting results for the entire sample and several sub-samples that are created by selecting the right tail of the empirical distribution at various percentiles. Head et al. (2014) note that the Pareto distribution is commonly used in the empirical trade literature because it is tractable, can be theoretically micro-founded, and empirically provides a good fit of the right tail of the observed distribution of manufacturing firm sales data. However, the left tail of this distribution is not as well approximated by the Pareto distribution, and it is not uncommon for researchers to estimate values of the shape parameter that violate the key assumption that was noted in Section 2.3, that $\theta > \sigma - 1$. Table 2.5 illustrates that this assumption is indeed violated for the full sample. Theoretically valid estimates of the shape parameter, θ , are obtained

when the sample is restricted to the top 50 percent or less (as ordered by size in value added). I follow Head et al. (2014) in reporting the welfare statistic for all theoretically valid estimates of θ .

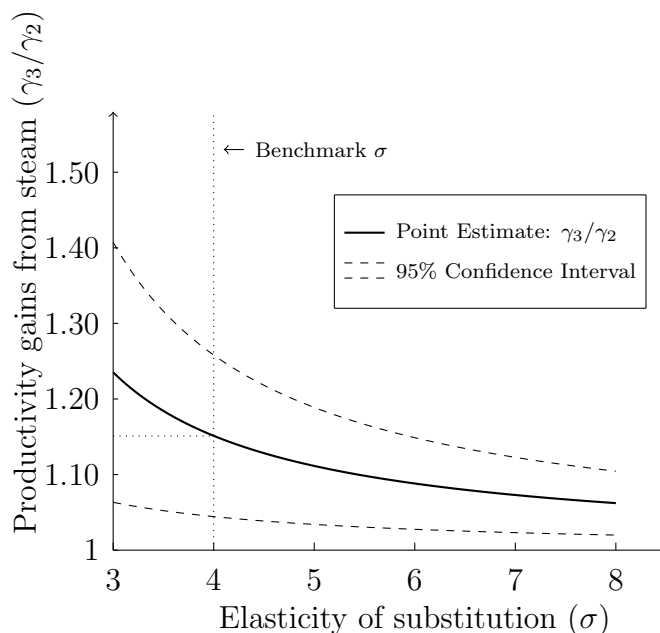


Figure 2.3: Firm-Level Productivity Gains from Steam Power Adoption

In Table 2.5 the estimates of the shape parameter, θ , use the parameterized value of $\sigma = 4$, and the estimated value of the parameter $-(\sigma - 1)/\theta$ from the QQ regression. The welfare statistic, \mathbb{W}/\mathbb{W}' , is then estimated using the equation (2.12) and the estimates of θ and the IV-WP estimate of the ratio $R/(R_1 + R_2 + R_{-3})$. The results of this calculation for the various estimates of θ are presented in Table 2.5. The results indicate that the introduction and adoption of steam power resulted in an increase in welfare that ranged from 3.0-5.2 percent depending on the estimate of θ that is used.

As previously noted, the framework developed in this paper draws on the literature on heterogeneous firms in international trade. The economic significance of the estimated welfare gains from steam power adoption can be appreciated by comparing the results in Table 2.5 with contemporary estimates of the welfare gains from trade.

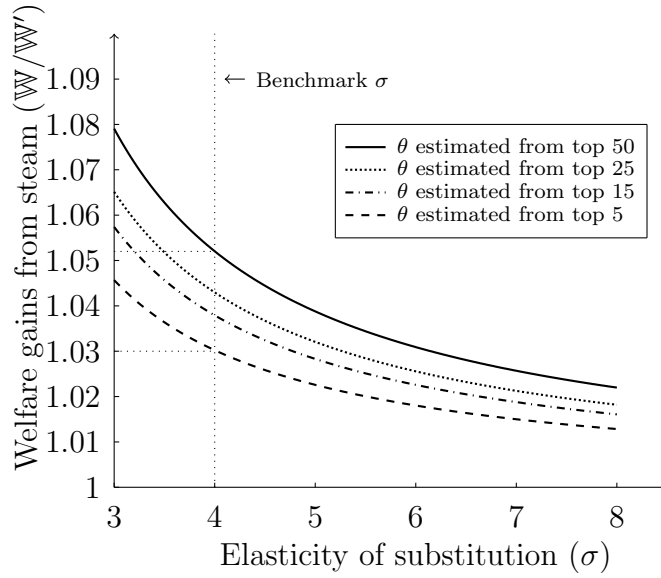


Figure 2.4: Aggregate Welfare Gains from Steam Power Adoption

For example, Melitz and Redding (2013) calibrate a model of heterogeneous firms in trade using moments from the U.S. data and estimate the welfare gains from opening a closed economy to trade. Their baseline calibration suggests that the welfare gains from trade are between 2 and 3 percent. Melitz and Redding (2013) use values of $\theta = 4.25$ and $\sigma = 4$ in their calibration. In Table 2.5, the estimated value of θ for the top 15% of firms is 4.184, which is close to the value used by Melitz and Redding (2013). Using this estimate of θ yields an estimated aggregate welfare gain of 3.8 percent from steam power adoption. Thus, the estimated welfare gains from steam power adoption are roughly one and a half times as large as contemporary estimates of the welfare gains from international trade.

The estimates of the productivity and welfare gains from steam power adoption are sensitive to the parameterized value of the elasticity of substitution, σ . Figures 2.3 and 2.4 explores the robustness of the baseline results by plotting the estimated productivity and welfare gains when alternative values of σ ranging from 3 to 8

are used. Figure 2.3 plots the point estimate and 95% confidence interval for the increase in firm-level productivity from firms switching from water to steam power. The estimated productivity gains are decreasing in σ and range from 23.5 to 6.2 percent. Figure 2.4 plots the estimated aggregate welfare gain from the introduction and adoption of steam power in the Canadian manufacturing sector in 1871. The various line plots correspond to the different sub-samples used in estimating the shape parameter θ . The estimated welfare gain from steam power adoption is also decreasing in σ and ranges from a high of 7.9 percent to a low of 1.3 percent.

2.5 Conclusion

In this paper I develop a tractable framework for calculating the firm-level productivity and aggregate welfare gains that are accrued from the introduction and adoption of a new technology. My analysis of steam power adoption in the late nineteenth century illustrates the potential of my model, and at the same time contributes to the literature on general purpose technologies. My results suggest that the introduction of a GPT can have significant impacts on aggregate welfare, even during the early stages of diffusion when only a small fraction of the economy has adopted the technology. Despite the fact that only 8 percent of Canadian manufacturing firms used steam in 1871, my results suggest that the adoption of steam power resulted in a 3.0-5.2 percent increase in welfare. This finding can be partially explained by the fact that firms using steam were large, having accounted for approximately 42 percent of Canadian manufacturing value added in 1870. A second factor explaining the large welfare gains is the estimated increase of 15.1 percent in firm-level productivity for firms using steam power relative to those using water power. Therefore, this paper finds that technological adoption by a relatively small number of firms can have a

pronounced effect on aggregate welfare when positive selection results in a reallocation of resources to the highly productive firms that are the first to make use of the technology.

The theoretical framework developed in this paper has a parsimonious structure, yet is able to replicate the empirical regularities of steam power adoption in nineteenth century. The methodology can also be extended to study other research questions pertaining to firm heterogeneity in technological adoption. For example, Ristuccia and Solomou (2014) find that the productivity benefits of GPT adoption vary across industries. The theoretical framework in this paper could be easily extended to estimate the sector-specific welfare gains from GPT adoption, in a similar fashion to the approach used in multi-sector models of international trade, such as Melitz and Redding (2014). The methodology in this paper can also be applied to study the welfare gains from other GPTs such as electricity and information and communications technologies (ICTs). Crafts (2004) finds that relative to ICTs, the contribution of steam technology to economic growth was modest. However the growth accounting framework used in Crafts (2004) ignores the productivity and welfare effects associated with firm heterogeneity. When these effects are incorporated in the theoretical framework, this paper finds that the welfare gains from steam power were roughly one and a half times as large as contemporary estimates of the welfare gains from trade. Research directed at measuring the productivity and welfare gains of steam relative to other GPTs is an important topic for future research.

Chapter 3: Between a Cap and a Higher Price: Modeling the Price of Dairy Quotas under Price Ceiling Legislation¹⁸

3.1 Introduction

After four decades, the Canadian dairy supply management system is burgeoning on crisis. The Trans-Pacific Partnership (TPP) negotiations and the Canada-European Union Comprehensive Economic and Trade Agreement (CETA) have increased pressure on Canada to reform its supply managed sectors of agriculture. As the industry faces trade liberalization, a recent OECD study highlights the poor economic performance of Canada's dairy sector relative to the dairy sector in eight other OECD countries, (Kimura and Thi, 2013). While the sector already appears ill-equipped to compete in international markets, recent domestic policy changes in the supply

¹⁸This chapter is a forthcoming article in the *Canadian Journal of Economics*, published by Wiley-Blackwell, Chernoff (2015). I acknowledge the *Canadian Journal of Economics* and Wiley-Blackwell for allowing this article to be included in my thesis as per the terms and conditions of the publisher's Copyright Assignment Form.

management system are further weakening the sector's competitive position.

To achieve the objective of providing a "fair" return to farmers,¹⁹ supply management relies on controlling output levels through the use of raw milk production quotas. The exchange of dairy quotas on provincial auctions encourages efficient allocation of dairy quotas across farmers. However, concern over the effect of rising quota prices on the cost of entry and expansion led Canada's two largest dairy producing provinces, Quebec and Ontario, to implement a cap on the price of quotas in 2007 and 2009 respectively. In Quebec, dairy quotas have traded at the price ceiling (currently at \$25,000 per unit) for over three years. With the cap in place, the demand for quotas has greatly exceeded supply, and the auction's ability to allocate quotas to the most efficient farmers has been significantly compromised.

The extraordinary price of dairy quotas is indicative of the magnitude of the loss in consumer surplus that arises from supply management. The price of dairy quotas reflects the profitability of dairy farming, which is inflated under supply management. In 2011, the value of raw milk was approximately 67% higher in Canada than the United States.²⁰ Higher farm prices are passed through the supply chain and result in higher consumer prices in Canada. For example, Larue (2012) estimates that average retail milk prices are 44% higher in Canada than in the United States. Higher retail prices result in a substantial reduction in consumer surplus as prior research indicates that consumers' own-price elasticity of demand is relatively inelastic across most dairy products.²¹ Furthermore, Larue (1994) illustrates that the system

¹⁹The *Canadian Dairy Commission Act* states that an objective of the Canadian Dairy Commission (CDC) is "to provide efficient producers of milk and cream with the opportunity of obtaining a fair return for their labour and investment..." (Canada, Department of Justice, 1985, s.8).

²⁰Author's own calculation using data sourced from the Canadian Dairy Information Centre and United States Department of Agriculture, Economic Research Service. Pounds of milk are converted to hectolitres using a conversion ratio of 2.273 lbs/litre. US dollars are converted to Canadian using the Bank of Canada's annual average exchange rate.

²¹For example, Moschini and Moro (1993) find that the own-price elasticities of milk, butter, cheese and other dairy products are -0.34, -0.92, -0.4, and -1.02 respectively.

is regressive by noting that per capita consumption of dairy products is relatively constant across low and high income households. Therefore supply management imposes a disproportionate burden on the income of lower income households.

In this paper I develop a theoretical model of supply management and use it to analyze several counterfactual experiments that are currently being discussed in the context of policy reform. I begin by modeling farmers' demand for dairy quotas and derive the equilibrium price contingent on the aggregate supply of quotas. The aggregate supply of quotas is endogenous to the model and is set by an administrator of the supply management system. The administrator has knowledge of farmers' cost of production and the demand for raw milk, and sets the aggregate supply of quotas to achieve a target level of profits for the average producer. I develop an econometric model to estimate a structural parameter of the model that is required for the counterfactual experiments. This parameter is estimated using farm-level data from the Quebec Federation of Management Clubs' Agritel database, and a quota price series from the Fédération des Producteurs de Lait du Québec (FPLQ).²²

Using my econometric results and Statistics Canada data on the average profit per unit of quota in Quebec, I estimate the equilibrium price of dairy quotas in Quebec for the period 1993 to 2011. The modeled price fits the actual market exchange price well during the pre-price ceiling era. During the years following the introduction of the price ceiling in 2007, the modeled price remains well above the price ceiling. If the price cap were removed, I estimate that dairy quotas in Quebec would trade at a price of \$29,923 per unit in 2011, which is 20% above the current ceiling price.

Cairns and Meilke (2012) argue that reducing the farm price of milk is a better means of reducing quota prices than price ceiling legislation. I estimate that lowering the valuation of quotas to \$25,000 per unit in 2011 would have required a 9.97%

²²The FPLQ is the provincial dairy marketing board in Quebec.

reduction in the farm price of milk. The theoretical model predicts that a decrease in the farm price of milk encourages the exit of small dairy farmers from the industry. I predict that removing the price ceiling and reducing the farm price of milk by 9.97% would result in a 34.6% increase in the minimum viable scale of production.

I consider two policy experiments that could result in a farm price reduction of the magnitude required to reduce quota prices to \$25,000 per unit.²³ The first policy experiment considers an expansion of the aggregate supply of quotas that is similar to the policy recommendations of Robson and Busby (2010). I predict that a 4.16% increase in production quotas would result in a 9.97% decline in the farm price and reduce the quota price to \$25,000 per unit. The second policy experiment considers a specific trade liberalization scenario that is comparable to a moderate outcome to the prospective TPP and CETA agreements. By combining my results with those of Abbassi et al. (2008), I find that a similar drop in the price of dairy quotas in Quebec could be achieved through a 25% reduction in bound tariffs and 50% increase in the MAC, assuming that the aggregate supply of raw milk in Quebec is held constant following trade liberalization.

The ascent of quota prices over the past two decades has sparked a small literature on Canadian dairy quota prices. Examples include Barichello (1996), Chen and Meilke (1998), Turvey et al. (2003), Doyon et al. (2006), and Cairns and Meilke (2012). Of the aforementioned studies, only Cairns and Meilke (2012) address the issue of price ceiling legislation. Cairns and Meilke (2012) use a capital asset pricing model to analyze the recent rise in quota prices, and assess the welfare effects for dairy farmers arising from price ceiling legislation.²⁴ Cairns and Meilke (2012) argue that

²³More generally, as Barichello et al. (2009) point out, there are several additional policy options in the broader context of reform of the supply management system.

²⁴Cairns and Meilke (2012) also analyze the welfare effects of the progressive transfer assessment policy that was introduced on the Ontario dairy quota exchange in 2006.

the rapid rise in the price of quotas during the late 1990s was the result of two factors. The first factor was a drop in farmers' perception of the risk of major reform to supply management following the conclusion of the Uruguay Round of multilateral trade negotiations in 1994. The second factor was an improvement in the borrowing conditions facing dairy farmers in the late 1990s. Based on 2007 data from the Ontario dairy quota exchange, Cairns and Meilke (2012) estimate that the deadweight loss from price ceiling legislation is \$4.1 million.

There are a number of features that distinguish this paper from Cairns and Meilke (2012) and other research on this topic. My theoretical model captures the interaction between farmers' optimal demand for quotas and the role of supply management in setting the aggregate supply of quotas. This feature enables counterfactual experiments that are not possible using the asset pricing model of Cairns and Meilke (2012), such as determining the price effect of expanding the aggregate supply of quotas. Another contribution of this paper are my empirical estimates of farmers' perception of policy risk. Cairns and Meilke (2012) use a calibration exercise and aggregate data from Ontario to estimate policy risk, whereas I estimate policy risk from farm-level data. My estimates of policy risk are similar to those of Cairns and Meilke (2012). By estimating policy risk from micro-level data, my results provide confirmation that the parameterization used in Cairns and Meilke (2012) is consistent with the behaviour of farms in my sample. Using a micro-founded model also allows me to complete farm-level analysis that is not possible using a capital asset pricing framework. In particular, I estimate the effect of a decrease in the farm price of milk on the exit decisions of small farmers. To the best of my knowledge, this paper is the first to establish linkages between the ceiling price of quotas, the farm price of milk, and trade liberalization. More generally, this model provides a new framework for evaluating future proposals to reform supply management, which is of particu-

lar relevance given the current policy environment. Finally, this paper contributes to the broader literature on price ceilings. The empirical studies in this literature focus mainly on cases where price ceilings have been applied to consumer goods and services.²⁵ In contrast, this paper analyzes the effect regulating the price of an input that constrains the entry and expansion of producers.

The remainder of the paper is organized as follows: Section 2 explains how production quotas are used as part of the supply management system in the Canadian dairy sector. Section 3 presents a model of dairy farmers' demand for production quotas. Section 4 presents the econometric model, empirical specification, and estimation results. Section 5 includes the counterfactual policy experiments, and Section 6 concludes.

3.2 Production Quotas, Supply Management, and the Canadian Dairy Sector

There are three main mechanisms used in the system of supply management in the Canadian dairy sector. These three mechanisms include administering the farm price of milk, restricting Canadian imports of dairy products, and setting milk production levels through the use of production quotas. The price Canadian dairy farmers receive for their milk is jointly administered by the Canadian Dairy Commission (CDC) and the provincial marketing boards. The CDC sets support prices for skim milk powder and butter. These support prices guide the provincial marketing boards in setting the prices that processors pay for raw milk destined for industrial purposes (butter, cheese, yogurt, etc.). The provincial marketing boards also set the price that

²⁵Recent examples in this literature include Knittel and Stango (2003) and Sen et al. (2011).

processors pay for “fluid” milk (table milk and cream). Farmers receive a “blended” price that reflects the various prices paid by processors for the different classes of fluid and industrial milk. Farm-level cost-of-production studies are used to guide the CDC and provincial marketing boards in setting prices at levels that provide a ‘fair’ return to farmers.

Production quotas are administered at the federal and provincial level, and are used to equate supply and demand at the administered price levels. The national marketing sharing quota (MSQ) determines production levels for industrial milk and is set at the federal level and then allocated to each province. In contrast, provincial production levels of fluid milk are determined by provincial marketing boards. Each provinces’ share of the MSQ, together with its provincial fluid milk production level, comprises its total production quota (TPQ). Dairy farmers buy and sell production quotas on provincial quota market exchanges, which are double auctions that operate on a monthly basis. Quotas are exchanged in units of one kilogram of butter fat per day and entitle a farmer to produce in perpetuity.²⁶ While quotas are bought and sold in units, they are properly understood as shares of the provincial TPQ as individual quota holdings are periodically adjusted on a percentage basis in step with changes to the TPQ.

Figure 3.5 illustrates that quota prices in Quebec tripled in value between the early 1990s and mid-2000s. A similar ascent in quota prices occurred in Ontario over this same time period. The escalating cost of financing dairy farm entry and expansion led the provincial marketing boards in Quebec and Ontario to introduce a price cap on quotas in 2007 and 2009 respectively. With the price cap limiting the price of quotas, all farmers with a reservation bid above the ceiling price bid for quotas, and there is

²⁶The annual production entitlement of one unit of quota is roughly equivalent to the milk produced by a single cow in one year.

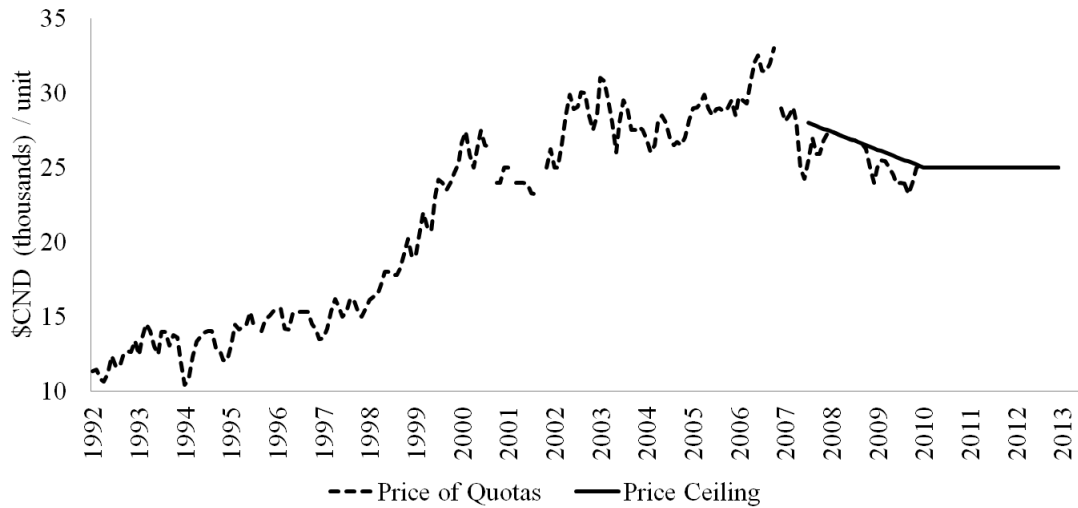


Figure 3.5: Quebec Dairy Production Quota Prices - 1993-2013

Source: FPLQ

no market-based mechanism for allocating quotas to the most efficient farmers. As a result, the FPLQ has established rationing rules for distributing available quotas to all farmers that are willing to pay the ceiling price. During the price ceiling era, the demand for quotas has vastly exceeded the quantity of quotas supplied to the quota exchange. For example, on the 2012 monthly quota exchange the quantity of quotas demanded at the price cap was approximately 20 times the quantity supplied. The magnitude of this imbalance between supply and demand is indicative of the inefficiency that has been created as a result of the price ceiling legislation.

3.3 A Model of Dairy Farmers' Demand for Quotas

In this section I develop a model of dairy farmers' demand for production quotas. Dairy farmers are assumed to be infinitely lived with preferences over a numeraire

consumption good. Farmers face a budget constraint and must allocate resources between units of consumption and units of quota. Each unit of quota entitles the farmer to one share of the TPQ. The timing of the model is such that farmer i enters period t with a quota holding of $q_{i,t}$. The government then announces the aggregate growth rate, g_t , of the TPQ, and farmer i is then entitled to produce $\tilde{q}_{i,t} \equiv (1 + g_t)q_{i,t}$ units in period t .

Each farmer produces milk using a technology that features constant returns to scale in production quotas. All farmers receive the government administered blended price for milk, P_t^m , but are heterogeneous with respect to their marginal cost of producing milk, $w_{i,t}$. For all farmers the random variable $w_{i,t}$ is assumed to be stationary with unconditional mean μ_w . The profit function of farmer i in period t is:

$$\Pi_{i,t} = P_t^m q_{i,t}(1 + g_t) - w_{i,t} q_{i,t}(1 + g_t) = \pi_{i,t} q_{i,t}(1 + g_t)$$

Where $\pi_{i,t} \equiv P_t^m - w_{i,t}$ is farmer i 's marginal profit at time t . After producing and collecting profits, the farmer must allocate resources between current consumption, $c_{i,t}$, and investment in quotas, $q_{i,t+1}$. When making this allocation decision the financial resources available to farmer i include earned income, $\pi_{i,t} q_{i,t}(1 + g_t)$, and asset wealth, $P_t q_{i,t}(1 + g_t)$; where P_t is the price of production quotas at time t .

At the end of each period there is a risk of the government eliminating supply management, a policy change that results in the profits from quotas falling to zero under the assumption of free entry of new farmers. As demonstrated below, in the period following such a policy reform the equilibrium price of quotas, and therefore the sum total of financial resources available to farmers, also drops to zero. To assist farmers in the transition to the new policy environment, I assume that the government awards each farmer with an adjustment payment that is dependent on the farmer's

quota holdings. Specifically, farmers are assumed to receive an adjustment payment equal to some fraction, η , of the market value of their quota holding. Following the policy reform, farmers earn an income y in each period through working in another sector of the economy.

Farmer i 's dynamic optimization problem can be framed recursively as follows:

$$V(q_{i,t}, \pi_{i,t}, g_t, P_t) = \max_{c_{i,t}, q_{i,t+1}} \ln(c_{i,t}) + \frac{1-\lambda}{1+r} E_t V(q_{i,t+1}, \pi_{i,t+1}, g_{t+1}, P_{t+1}) + \frac{\lambda}{1+r} \left(\ln(\eta P_t q_{i,t+1}) + \frac{\ln(y)}{r} \right) \quad (3.20)$$

$$\text{subject to : } P_t q_{i,t}(1 + g_t) + \pi_{i,t} q_{i,t}(1 + g_t) = c_{i,t} + P_t q_{i,t+1}$$

$V(q_{i,t}, \pi_{i,t}, g_t, P_t)$ is the value of a dairy farm with quota holding $q_{i,t}$ and marginal profit $\pi_{i,t}$, when the growth rate of quotas is g_t and the price of quotas is P_t . The parameter r is the discount rate, and λ is farmers' perception of the probability of policy reform. The left-hand side of the budget constraint represents the financial resources available to the farmer, and the right-hand side represents expenditures.

Define $c_{i,t}^*(q_{i,t}, \pi_{i,t}, g_t, P_t) \equiv c_{i,t}^*$ and $q_{i,t+1}^*(q_{i,t}, \pi_{i,t}, g_t, P_t) \equiv q_{i,t+1}^*$ as the policy functions for consumption and quotas respectively that solve (3.20). Under the specification of log utility from consumption, the policy functions $c_{i,t}^*$ and $q_{i,t+1}^*$ have the following closed form solutions²⁷:

$$c_{i,t}^* = \frac{r + \lambda}{1 + r + \lambda} (P_t + \pi_{i,t}) q_{i,t}(1 + g_t) \quad (3.21)$$

$$P_t q_{i,t+1}^* = \frac{1}{1 + r + \lambda} (P_t + \pi_{i,t}) q_{i,t}(1 + g_t) \quad (3.22)$$

The policy functions have an intuitive interpretation. Each period, farmer i allocates a percentage, $(r + \lambda)/(1 + r + \lambda)$, of available resources to consumption, and a percentage, $1/(1 + r + \lambda)$, to quotas. If policy risk is high then farmers will purchase

²⁷See Appendix B for a detailed derivation of the policy functions $c_{i,t}^*$ and $q_{i,t+1}^*$.

less quota since there is a lower probability of capturing future rents from quotas.

Dairy farmers are assumed to take the price of quotas, P_t , as given. Dividing equation (3.22) by P_t and summing across all n farmers yields the aggregate demand for quotas, Q_{t+1}^d . The aggregate supply of quotas, $\tilde{Q}_t = Q_t(1 + g_t)$, is the beginning of period TPQ adjusted by the aggregate growth rate g_t . Equating aggregate supply and demand, $Q_{t+1}^d = \tilde{Q}_t$, yields a parsimonious pricing formula for quotas:

$$\begin{aligned}
Q_{t+1}^d &= \sum_{i=1}^n q_{i,t+1}^* = \sum_{i=1}^n \left[\left(\frac{1}{1+r+\lambda} \right) q_{i,t}(1+g_t) + \left(\frac{1}{1+r+\lambda} \right) \frac{\pi_{i,t}}{P_t} q_{i,t}(1+g_t) \right] \\
&= Q_t(1+g_t) = \tilde{Q}_t \\
\frac{1}{1+r+\lambda} \sum_{i=1}^n q_{i,t}(1+g_t) + \frac{1}{1+r+\lambda} \frac{1}{P_t} \sum_{i=1}^n \pi_{i,t} q_{i,t}(1+g_t) &= Q_t(1+g_t) \\
P_t &= \frac{\frac{1}{1+r+\lambda}}{\left(1 - \frac{1}{1+r+\lambda}\right)} \frac{\sum_{i=1}^n \Pi_{i,t}}{Q_t(1+g_t)} = \frac{\bar{\pi}_t}{r+\lambda}, \tag{3.23}
\end{aligned}$$

where $\bar{\pi}_t \equiv \sum_{i=1}^n \Pi_{i,t} / (Q_t(1+g_t))$ is the average profit per unit of quota at time t . Note that when policy risk is zero, $\lambda = 0$, the price reduces to the familiar asset pricing formula $P_t = \bar{\pi}_t / r$. More generally, the price of quotas is decreasing in policy risk and the discount rate, and increasing in average profit per unit of quota. In the event of policy reform quota rents fall to zero, $\pi_{i,t} = 0 \ \forall i$, and it follows immediately from (3.23) that the price of quotas falls to zero.

The role of supply management in determining the growth rate of the TPQ, g_t , can be incorporated in this model in a manner that is consistent with the objectives of the system.²⁸ At the beginning of each period an administrator of the system conducts a cost-of-production study after each farmer has observed their marginal

²⁸I am grateful to an anonymous referee who recommended that I explicitly model the role of supply management and the linkage between the growth rate of quotas, g_t , the farm price of milk, P_t^m .

cost of production for the period $w_{i,t}$. The administrator also knows the quota holding of each producer, the beginning of period TPQ, and the aggregate demand for raw milk as a function of the farm price.²⁹ Based on this information, the administrator chooses g_t so that the average profit per unit of quota remains constant at $\bar{\pi}$, which is exogenously chosen by the system in accord with its objective to provide ‘fair’ return to farmers. Therefore, the administrator’s objective is to choose g_t each period to satisfy the following constraint:

$$\bar{\pi} = \frac{\sum_{i=1}^n \Pi_{i,t}}{Q_t(1 + g_t)} = P_t^m(Q_t, g_t) - \frac{\sum_{i=1}^n w_{i,t}q_{i,t}}{Q_t} \quad (3.24)$$

There is a unique g_t that solves the administrator’s problem under the assumption that the demand for raw milk is strictly decreasing in the farm price. Note that in this modeled policy environment the price of quotas is constant, $P = \bar{\pi}/(r + \lambda)$. Variations in the modeled price of quotas are the result of changes in the discount rate, the policy risk parameter, or changes in the targeted average profit per unit of quota.

I now consider the case of a deterministic equilibrium where each farmer’s marginal cost is fixed at the unconditional mean μ_w . In this equilibrium it is possible to characterize the size profile of farms that exit the dairy industry. Starting from equation (3.24) it follows that the deterministic steady state farm price is constant $P_{ss}^m = \bar{\pi} + \mu_w$, and hence $g_{ss} = 0$. As $g_{ss} = 0$, the TPQ is held constant at a level Q_{ss} that supports the equilibrium farm price. g_{ss} and P can be substituted into equation (3.22) to solve for farmer i ’s equilibrium quota holding:

$$q_{i,ss} = \frac{1}{1 + r + \lambda} \left(1 + \frac{\bar{\pi}}{P}\right) q_{i,ss} = q_{i,ss}$$

²⁹Recall that the aggregate supply of raw milk is simply $\tilde{Q}_t = Q_t(1 + g_t)$.

Hence each farmer's output level is constant in equilibrium and *any* distribution of quotas across farmers is an equilibrium provided that it satisfies the aggregation condition: $\sum_{i=1}^n q_{i,ss} = Q_{ss}$.

The steady state value function of a farmer with quota holding $q_{i,ss}$ is:³⁰

$$V(q_{i,ss}) = \frac{\lambda}{r + \lambda} \left(\ln \left(\frac{\eta}{r + \lambda} \right) + \frac{1}{r} \ln(y) \right) + \frac{1 + r + \lambda}{r + \lambda} \ln(\bar{\pi} q_{i,ss}) \quad (3.25)$$

Farmers that choose to exit the dairy industry do so by selling their entire stock of quotas in the quota exchange after production occurs in the period. After exiting a farmer earns an income y in all future periods by working in another sector of the economy. The value of a farmer that exits with quota holding $q_{i,ss}$ is therefore $V^e(q_{i,ss}) = \ln((P + \bar{\pi})q_{i,ss}) + \ln(y)/r$. A farmer will exit the industry if $V(q_{i,ss}) < V^e(q_{i,ss})$. As $\partial V(q_{i,ss})/\partial q_{i,ss} > \partial V^e(q_{i,ss})/\partial q_{i,ss}$, farmers with small quota holdings are more likely to exit the industry than are large farmers. By equating the two value functions it is possible to derive the unique quota level q_{ss}^* at which a farmer is indifferent between exiting and remaining in the industry.

$$q_{ss}^* = \frac{y(1 + r + \lambda)^{(r+\lambda)}}{\bar{\pi}\eta^\lambda(r + \lambda)^r} \quad (3.26)$$

Since all farmer's with quota holdings less than q_{ss}^* exit the industry, q_{ss}^* is the minimum viable scale of production in the dairy industry for the deterministic equilibrium. The minimum scale of production is decreasing in the average profit per unit of quota, $\bar{\pi}$. Intuitively, when profits are high small farmers are less likely to exit. The minimum scale of production is also decreasing in η , which implies that small farmers are less likely to exit as the adjustment package becomes more generous. The effect of a

³⁰See Appendix B for the derivation of the value function, $V(q_{i,ss})$.

marginal increase in the policy risk parameter, λ , on q_{ss}^* is indeterminate of sign.

In the Appendix B I illustrate that it is possible to analyze the effects of price ceiling legislation on the minimum viable scale of production. Under price ceiling legislation, the minimum viable scale of production is:

$$q_{ss}^* = y\eta^{-\lambda}(P + \bar{\pi})^{-\frac{(1-\lambda)(1+2(r+\lambda))}{r+\lambda}} P^{\frac{1+r-2\lambda(r+\lambda)}{r+\lambda}} (r + \lambda)^{-(1+r)}(1 + r + \lambda)^{\frac{(1+r)(r+\lambda)+(1-\lambda)(1+r+\lambda)}{r+\lambda}} \quad (3.27)$$

In the neighborhood of the equilibrium price of quotas, decreasing quota prices by price ceiling legislation may either decrease or increase the minimum viable scale of production, depending on the values of the discount rate, r , and the policy risk parameter, λ . In section 4.2 I estimate that farmers' perception of policy risk was 4.63% in the period preceding the implementation of the price ceiling in 2007. In that section I follow Cairns and Meilke (2012) in assuming that farmers' discount rate during this period was equal to the prime interest rate plus two percent, 7.262%. When evaluated at these estimates, decreasing the quota prices through price ceiling legislation *reduces* the minimum viable scale of production. This result suggests that the implementation of price ceiling legislation in 2007 effectively discouraged small farmers from exiting the industry in Quebec.

3.4 Empirical Application

3.4.1 Econometric Model and Variable Specification

In this section I develop an econometric model to estimate the structural parameter $1/(1 + r + \lambda)$, which is required for the counterfactual policy experiments in section 5. Specifically, I divide both sides of equation (3.22) by financial resources, $P_t q_{i,t}(1 +$

$g_t) + \pi_{i,t}q_{i,t}(1 + g_t)$, to isolate this parameter. Adding an error term to account for unobserved heterogeneity results in the following empirical specification:

$$\frac{P_t q_{i,t+1}}{(P_t q_{i,t}(1 + g_t) + \pi_{i,t} q_{i,t}(1 + g_t))} = \theta_1 + u_{i,t} \quad (3.28)$$

Under the assumption $E(u_{i,t}) = 0$, equation (3.28) implies that farmers' mean percentage allocation of resources to quota purchases equals $\theta_1 = 1/(1 + r + \lambda)$.

Previous research suggests that the discount rate r and policy risk λ were not constant over the period 1993-2011. In particular, Cairns and Meilke (2012) argue that policy risk declined following the conclusion of the Uruguay Round of multilateral trade negotiations in 1994. Cairns and Meilke (2012) also note that the borrowing conditions facing dairy farmers improved in the late 1990s. While I have not included a credit constraint in my model, a simple means of incorporating changes in the borrowing environment is to allow the interest rate, r , to vary. Specification (3.28) can be easily modified to allow for structural breaks in the parameter $1/(1 + r + \lambda)$ by introducing break year indicator variables. In particular, I allow for up to two structural breaks in the years t_{b1} and t_{b2} as follows:

$$\frac{P_t q_{i,t+1}}{(P_t q_{i,t}(1 + g_t) + \pi_{i,t} q_{i,t}(1 + g_t))} = \theta_1 + \theta_2 d_{1,t} + \theta_3 d_{2,t} + u_{i,t} \quad (3.29)$$

$$d_{1,t} = \begin{cases} 1 & \text{if } t \leq t_{b1} \\ 0 & \text{otherwise} \end{cases} \quad d_{2,t} = \begin{cases} 1 & \text{if } t_{b1} < t \leq t_{b2} \\ 0 & \text{otherwise} \end{cases}$$

Under this specification the value of $1/(1 + r + \lambda)$ is $\theta_1 + \theta_2$ between t_0 and t_{b1} ; equal to $\theta_1 + \theta_3$ between $t_{b1} + 1$ and t_{b2} ; and equal to θ_1 in the years thereafter. The timing of the break years is determined by an econometric specification test described in the

estimation results section of the paper.

The dependent variable in equation (3.29) includes both aggregate variables, P_t and g_t , and farm-level variables, $q_{i,t}$ and $\pi_{i,t}$, that must be specified. The price of quotas, P_t , is specified as the annual average price of quotas (dollars per hectolitre) on the monthly dairy quota exchange in Quebec.³¹ As a proxy for the growth of Quebec's TPQ, g_t , I use the percentage change in Quebec's share of the national MSQ. Ideally I would use the percentage change in Quebec's TPQ, however prior to 2002 data on Quebec's TPQ are only available for the years 1996, 1997, and 2003. Quebec's share of the national MSQ is sourced from the Canadian Dairy Information Centre where it is reported as of August 1st of each year.

The farm-level variables are specified using the Agritel database, a farm-level panel data set that is collected annually by the Federation of Management Clubs in the province of Quebec. Under the specification of constant returns to scale in quotas, the marginal profit per unit of quota, $\pi_{i,t}$, is equal to the average profit per unit of quota. Therefore the net operating profit of a farmer (total revenue net of variable costs) divided by the farmer's quota holding is an appropriate measure for marginal profits. Finding an accurate measure of the variable costs associated with dairy production poses difficulties given the data that are available. The challenge arises because many dairy farmers are engaged in *joint production* of various agricultural outputs. For example a dairy farmer may produce field crops and use a portion of the yield as feed for dairy cattle and then sell the remaining crops. Failing to account for the non-dairy costs in variable costs could result in measurement error and biased estimates. To address this problem I use a revenue threshold to eliminate farmers that are heavily involved in joint production of agricultural outputs. Specifically,

³¹To convert the price of dairy quotas to dollars per hectolitre, I use milk composition annual average test data for the province of Quebec sourced from the Canadian Dairy Information Centre.

I include only those farmers whose annual income from milk production and dairy subsidies account for at least 80% of total revenue. The use of revenue thresholds to address the problem of joint agricultural production is a common methodology in economic studies of the dairy sector.³² After applying the revenue threshold to filter the data, the marginal profit variable, $\pi_{i,t}$, is defined as total revenue net of variable costs divided by the farmer's quota holding.³³

Specifying the quota holding variable requires careful consideration of differences between the latent specification of the model and the observed values in the data. Recall that the definition of $q_{i,t}$ is the beginning of period t quota holding of farmer i . Also recall that $\tilde{q}_{i,t} \equiv q_{i,t}(1 + g_t)$ is defined as the end of period t quota holding of farmer i . In my empirical specification I proxy a farmer's end of period quota holding with their annual milk production (in hectolitres).³⁴ To accommodate for the fact that I proxy for farmer i 's end of period quota holding, specification (3.29) needs to be modified. Specifically, by multiplying the dependent variable by $(1 + g_{t+1})/(1 + g_t)$ and using the definition of $\tilde{q}_{i,t}$, equation (3.29) can be re-written:

$$\frac{P_t \tilde{q}_{i,t+1}}{(P_t \tilde{q}_{i,t}(1 + g_{t+1}) + \pi_{i,t} \tilde{q}_{i,t}(1 + g_{t+1}))} = \theta_1 + \theta_2 d_{1,t} + \theta_3 d_{2,t} + u_{i,t} \quad (3.30)$$

Where $d_{1,t}$ and $d_{2,t}$ are defined as in equation (3.29). Note that the only difference between equation (3.29) and equation (3.30) is that $\tilde{q}_{i,t}$ and g_{t+1} are substituted in place of $q_{i,t}$ and g_t respectively.

³²Alvarez et al. (2006), Godah et al. (1990), and Boots et al. (1997) use revenue thresholds of 90, 70, and 50 percent respectively.

³³As an additional measure of safety against measurement error I also omit suspected outliers. In each period t , suspected outliers for the variable $\pi_{i,t}$ are identified using the standard 1.5 interquartile range (IQR) rule. Specifically, I identify suspected outliers as those values of $\pi_{i,t}$ that fall a distance of more than 1.5xIQR below the first quartile in period t , or 1.5xIQR above the third quartile in period t .

³⁴This is a reasonable proxy given that the FPLQ enforces strict financial penalties against dairy farmers producing under or over the amount permitted by their quota holding.

Over the observation period 1993-2011 the Agritel database represents between 5.05% and 9.06% of the population of Quebec dairy farmers. Comparing the average revenue of dairy farmers in the database to the population of Quebec dairy farmers reveals that farms in the Agritel database tend to be larger than the average dairy farm in Quebec. Over the period 1993-2011, average revenue for farms in my sample are 42% higher than for the population of dairy farms in Quebec. As the Agritel database is not a representative sample, it is important to question how sampling bias might affect the estimation results. With respect to the policy risk parameter, λ , it is possible that large farmers might perceive a lower level of policy risk if greater resources provide these farmers with better information and the ability to influence policy. However, given that in 2011 even the largest farmer in the Agritel sample produced only 0.25% of the TPQ for Quebec, it is reasonable to assume that the degree of influence of any one farmer is relatively small and therefore this effect is unlikely to bias the estimate of the parameter $1/(1 + r + \lambda)$.

Table 3.6 lists estimates of the average profit per unit of quota in Quebec for the years 1993-2011. The provincial profit estimates are calculated using the detailed revenues and costs data that are published in Statistics Canada's Whole Farm Database (WFD). Statistics Canada's WFD estimates revenues and costs at the provincial level using samples of individual farm tax records. An advantage of the WFD is that cross-tabulations are available by farm type. For example, a farm is classified as a dairy farm if "40% or more of agricultural sales are derived from the sale of dairy products and 10% or more from raising and selling dairy cattle" (p. 23, Statistics Canada, 2011). Revenue from dairy products and subsidies is included in the WFD and can be used to estimate dairy revenues. Certain categories of operating costs listed in the WFD are exclusive (or nearly exclusive) to dairy farming. However, other categories of operating costs are shared between dairy farming and other agricultural

enterprises. For the latter category I introduce a weight to allocate a share of these costs to dairy farming. Specifically, for each year the fraction of joint operating costs allocated to dairy farming is set equal to the fraction of total operating revenues derived from dairy products and subsidies. For the years 1993-2011 the average value of this ratio is 0.78. The procedure used to estimate the average profit per unit of quota in Quebec in each year is summarized in the following formula:

$$\frac{\Pi_t}{\tilde{Q}_t} = \frac{R_t - C_{1,t} - \nu_t C_{2,t}}{\tilde{Q}_t} = \frac{R_t - C_t}{\tilde{Q}_t} \quad (3.31)$$

The variables in equation (3.31) are defined as follows: R_t is dairy product and subsidy revenue in year t ; $C_{1,t}$ are operating costs that are specific to dairy farming in year t (feed, supplement, straw, and bedding expenses; veterinary fees, medicine and breeding fees; other livestock expenses; marketing expenses); $C_{2,t}$ are operating costs that are jointly shared between dairy and other agricultural enterprises in year t (crop expenses; farm share of fuel, machinery, truck, and automobile expenses; land rental expenses; custom work and machine rental expenses); ν_t is the period t weight introduced to account for joint agricultural production, $\nu_t = R_t/TR_t$, where TR_t is total operating revenues in year t ; C_t are operating costs in year t , $C_t = C_{1,t} + \nu_t C_{2,t}$; and \tilde{Q}_t is total annual milk production in year t (in hectolitres).³⁵

Table 3.6 presents data on the Agritel sample and the Quebec and New York Dairy farming industries. The data from New York are presented to contrast trends that are industry specific from those that are specific to the Quebec dairy sector and supply management.³⁶ Over the period 1993-2011 the number of dairy farms declined by 51.5% in Quebec and 53.18% in New York. The similarity of this trend across

³⁵The variable \tilde{Q}_t is sourced from Statistics Canada CANSIM series v382194.

³⁶In 2011, New York State was the 4th largest dairy producing state in the United States. As a large dairy producing state in a similar physical environment, the New York dairy farming industry provides a natural benchmark for comparison.

Table 3.6: Summary Statistics for Agritel Database, Quebec, and New York Dairy Farming

Year	Agritel		Quebec			New York
	Sample Size	Revenue	Population	Revenue	Π_t/\tilde{Q}_t	Population
1993	606	\$154,188	11,990	\$101,814	\$26.97	11,000
1994	649	\$167,803	11,490	\$112,902	\$27.83	10,700
1995	696	\$175,660	11,240	\$125,596	\$29.23	10,000
1996	624	\$177,949	10,200	\$131,192	\$27.35	9,200
1997	555	\$189,972	9,590	\$138,791	\$26.03	9,000
1998	531	\$199,417	9,385	\$146,395	\$27.51	8,700
1999	485	\$209,428	8,840	\$158,355	\$28.60	8,200
2000	459	\$242,518	8,860	\$173,633	\$31.79	7,900
2001	458	\$262,323	8,650	\$185,067	\$31.94	7,300
2002	462	\$275,490	8,120	\$193,412	\$32.51	7,200
2003	453	\$308,745	7,710	\$216,117	\$33.19	7,100
2004	420	\$341,000	7,375	\$231,020	\$33.18	6,900
2005	492	\$359,763	7,165	\$255,051	\$36.75	6,700
2006	484	\$385,434	6,850	\$269,088	\$37.62	6,400
2007	453	\$415,681	6,550	\$292,017	\$38.70	5,700
2008	465	\$429,115	6,300	\$311,945	\$39.17	
2009	542	\$476,373	6,190	\$324,565	\$37.44	
2010	527	\$495,613	6,005	\$342,534	\$39.50	5,250
2011	527	\$525,319	5,815	\$359,004	\$39.55	5,150

Data sources: Sample size and revenue - Agritel database. Population of dairy farmers, revenue, and Π_t/\tilde{Q}_t for Quebec: Statistics Canada's Whole Farm Database and CANSIM series v382194 and v31191101. Population of farms for New York: United States Department of Agriculture, Economic Research Service. Revenue is defined as mean revenue from dairy products and subsidies. Π_t/\tilde{Q}_t is defined as the average profit per hectolitre.

the two jurisdictions suggests that net exit is a feature of the dairy industry and is being driven by factors other than supply management. During the late 1990s there was a decline in the rate of net exit of Quebec dairy farmers. In particular from 1999 to 2000 the number of farms actually increased, and this coincides with the year of highest growth in the average profit per unit of quota (11.15%). This suggests the rising profit levels may have discouraged dairy farmers in Quebec from exiting during this period.³⁷ The monotonic increase in the average revenue per farm in Quebec from 1993-2011 can be explained by the steady decline in the number of farms and rise of aggregate revenue over this period.³⁸ A final observation from Table 3.6 concerns the average profit per unit of quota in Quebec, which grew at an annual average rate of 2.25% over the period 1993-2011. This suggests that the rapid increase in the price of quotas over the past two decades can be partially explained by a growth in profitability of dairy farming in Quebec.

3.4.2 Estimation Results

Model (3.30) is estimated over the pre-price ceiling era 1993-2005 with an unbalanced panel of 6,890 observations. To determine the preferred specification of the break years, model (3.30) is estimated by fixed effects using ninety-one different specifications. Seventy-eight of the specifications allow for two structural breaks in the parameter $1/(1 + r + \lambda)$ in every possible combination of years between 1993 and 2005. Twelve of the specifications allow for a single structural break in this parameter for each respective year between 1993 and 2004. The final specification estimates

³⁷Anecdotal evidence suggests that there are few dairy farmers entering the dairy industry relative to the number exiting. For example, in an effort to stimulate entry the marketing boards in several provinces (including Quebec and Ontario) have implemented ‘new entrant programs’ that subsidize the quota purchases of new farmers.

³⁸Aggregate revenue from dairy products and subsidies grew at an annual average rate of 3.09% per year over the period 1993-2011.

the model without any structural breaks. I then use the Akaike information criterion (AIC) and Bayesian information criterion (BIC) as the basis for model selection across the ninety-one different specifications. Both AIC and BIC select the model with structural breaks in the years 1995 and 1998.

Table 3.7 presents the fixed effects and random effects estimates of model (3.30) under the specification of heteroskedastic robust standard errors.³⁹ A test of overidentifying restrictions following the approach suggested by Arellano (1993) is used to choose between the random effects and fixed effects estimates. The fixed effects estimates are selected as preferred to the random effects estimates based on the results of this test.⁴⁰

Table 3.7: Fixed Effects and Random Effects Estimation Results

Coefficients	Random Effects	Fixed Effects
θ_1	0.895*** (0.00123)	0.894*** (0.00103)
θ_2	-0.0771*** (0.00180)	-0.0730*** (0.00227)
θ_3	-0.0309*** (0.00193)	-0.0286*** (0.00220)
Break years	1995, 1998	1995, 1998
R-squared		0.122
Observations	6,890	6,890
Number of producers	1,616	1,616

Preferred specification in bold; Robust standard errors in parentheses; *** p<0.01

For the purposes of modeling the price of quotas, the parameter of interest is $1/(1+r+\lambda)$. For the years 1993-1995 this parameter is estimated by $\theta_1 + \theta_2$, for 1996-1998 it is estimated by $\theta_1 + \theta_3$, and finally for 1999-2005 the parameter is estimated

³⁹The fixed effects estimates are calculated using the fixed effects (within) regression estimator. Fixed effects specified at the farm level. The random effects estimates are calculated using the generalized least squares random effects estimator.

⁴⁰The p-value of the test is 0.04.

by θ_1 . Under the preferred specification, $1/(1+r+\lambda) = 0.821$ for 1993-1995, $1/(1+r+\lambda) = 0.865$ for 1996-1998, and $1/(1+r+\lambda) = 0.894$ for 1999-2005. The timing of the first increase in the parameter of interest coincides with the conclusion of the Uruguay Round of negotiations at the end of 1994. This suggests that this increase may have been the result of a decline in dairy farmers' perception of the risk of major policy reform to supply management. The second structural break in the parameter of interest occurred in 1998. The timing and direction of this break is consistent with the theory that improved borrowing conditions in the mid-to-late 1990s led to increased demand and higher prices in Canadian dairy quota markets. However, as I have not explicitly modeled the credit constraint facing farmers, it is not possible to interpret the magnitude of this shift in relation to the borrowing and lending practices of dairy farmers.

The parameters r and λ are not individually identified through estimating equation (3.30). However, it is possible to estimate the change in perceived policy risk by assuming that farmers' discount rate was equal to the average prime interest rate plus two percent over the three sub-periods, $r_{93-95} = 0.09153$, $r_{96-98} = 0.07875$, and $r_{99-05} = 0.07262$. Substituting these interest rates into the parameter of interest yields estimates of policy risk of $\lambda_{93-95} = 12.69\%$, $\lambda_{96-98} = 7.71\%$, and $\lambda_{99-05} = 4.63\%$ in the three sub-periods respectively.⁴¹ My estimates of policy risk are close to those of Cairns and Meilke (2012), who estimate that policy risk fell from 10.7% in 1995 to 5.2% in 2006.

In sum, the econometric results provide evidence in support of the theory that reduced policy risk and improved borrowing conditions in the 1990s put upward pressure on the price of dairy quotas in Quebec. However, as noted in the previous sub-section, quota prices may have also risen due to increases in the average profit

⁴¹The changes in policy risk in 1996 and 1998 are statistically significant at the 1% level.

per unit of quota over the past twenty years. In the next section I combine these two factors by modeling price of dairy quotas over the period 1993-2011.

3.5 Counterfactual Policy Experiments

3.5.1 The Price of Dairy Quotas in the Absence of Price Ceiling Legislation

In this section I model the price of quotas over the period 1993-2011 using the formula for the equilibrium price of quotas from the model, $P = [1/(1 + r + \lambda)] / [1 - 1/(1 + r + \lambda)] \bar{\pi}_t$. There are two factors in the modeled quota price. The first term involves only the parameter $1/(1 + r + \lambda)$, and the second term is the average profit per unit of quota in each period. Average profit per unit of quota is estimated from Statistics Canada under the specification defined in equation (3.31).⁴² The parameter $1/(1 + r + \lambda)$ is specified over four sub-periods using the preferred estimates from Table 3.7. For the sub-periods 1993-1995, 1996-1998, and 1999-2005 I use the Table 3.7 estimates for each respective sub-period. For the final sub-period, 2006-2011, I use the estimate of $1/(1 + r + \lambda)$ for the period 1999-2005 ($\theta_1 = 0.894$). Implicit in this specification is the assumption that policy risk, λ , and the interest rate, r , did not change from their values during the period 1999-2005. This is clearly not the case as interest rates have fallen, while policy risk has likely risen as the TPP and CETA trade negotiations have increased the probability of supply management being reformed. Note that while lower interest rates increase the parameter $1/(1 + r + \lambda)$, higher policy risk decreases this parameter. I proceed under the assumption that the increase in λ has offset the decrease in r , and therefore the estimate of the parameter $1/(1 + r + \lambda)$ for

⁴²Milk production output is converted from hectolitres per year to kilograms of butterfat per day using a province specific conversion factor sourced from the Canadian Dairy Information Centre.

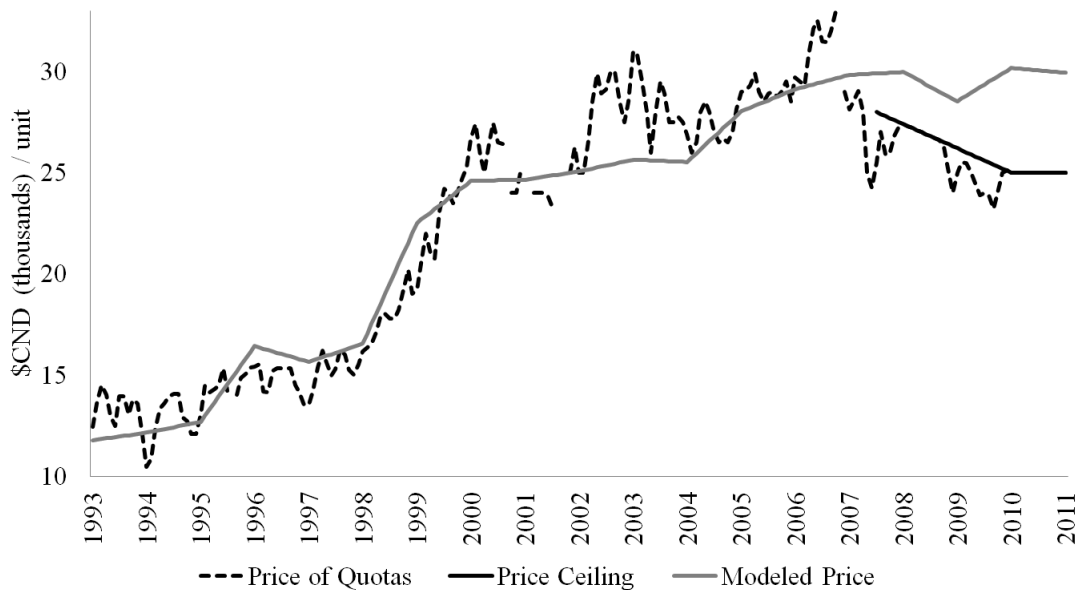


Figure 3.6: Quebec Dairy Quota Prices: Exchange, Ceiling, and Modeled

the period 1999-2005 provides a reasonable estimate for the period 2006-2011.

A second assumption implicit in my methodology concerns the effect of the price ceiling on the average profit per unit of quota. It is likely that average production costs would be lower if the price ceiling were removed.⁴³ I assume that the decrease in average production costs would be offset by a reduction in the administered farm price of milk so as to hold the average profit per unit of quota constant. This assumption is reasonable given that supply management employs cost-of-production pricing to target a profit level that is set by the administrators of the system. Therefore, under the theoretical assumptions of the system and my model, changes in the average cost of production arising from the price ceiling affects the farm price but not price of quotas.

The modeled price series is presented in Figure 3.6, along with the actual price

⁴³As discussed in section 2, during the price ceiling era the FPLQ has implemented a rationing rule to allocate the available quotas to all farmers bidding at least the price cap. As a result, average production costs are higher under the price ceiling because there are fewer quotas in the hands of the most efficient producers than under the market allocation.

series, and the price ceiling. As Figure 3.6 illustrates, the modeled price series fits the actual price series relatively well throughout the pre-price ceiling era 1993-2005. In particular, the modeled price series closely matches the abrupt increase in the price of dairy quotas during the 1990s. The success of the model in this regard can be attributed to the econometric estimates of the timing and magnitude of the structural breaks in the parameter $1/(1+r+\lambda)$. While the two upward shifts in this parameter partially explain the rising trajectory of the modeled price, growth in the average profit per unit of quota is a second factor influencing this trend. The fact that the modeled price closely follows the actual prices series suggests that these two factors are important determinants of quota price dynamics during this era. In the period following the introduction of the price ceiling the modeled price of quotas remains well above the price cap. In 2011, I estimate that dairy quotas in Quebec would have traded at a price of \$29,923 per unit, which is 20% above the price ceiling of \$25,000.

3.5.2 Reducing Quota Prices by Lowering the Farm Price of Milk

In this section I estimate the magnitude of the decrease in the farm price of milk required to reduce the valuation of Quebec dairy quotas to the current ceiling price of \$25,000 per unit. Re-arranging equation (3.23) and using the definition of average farmer profits ($\pi_t = P_t^m - \sum_{i=1}^n w_{i,t}q_{i,t}/Q_t$) from equation (3.24) yields:

$$P_t^m = \left[\frac{1 - \frac{1}{1+r+\lambda}}{\frac{1}{1+r+\lambda}} \right] P_t + \frac{\sum_{i=1}^n w_{i,t}q_{i,t}}{Q_t} \quad (3.32)$$

Define P^* as the target price of quotas and $\$P_t^{m*}$ as the farm price of milk that supports this target. It is possible to solve for the value of $\$P_t^{m*}$ that supports a

target quota price of \$25,000 per unit given estimates of average operating costs $C_t/\tilde{Q}_t = \sum_{i=1}^n w_{i,t}q_{i,t}/Q_t$, and the parameter $1/(1+r+\lambda)$. The estimated values of \tilde{Q}_t and C_t are defined in equation (3.31), and $1/(1+r+\lambda)$ is set equal to its econometrically estimated value for the period 1999-2005. An important caveat to this methodology is the point raised in section 5.1 regarding the effect of the price ceiling on the farm price of milk. As argued in that section, if the price ceiling were removed the average cost of production would fall, as would the farm price, assuming that the administrators of supply management remain committed to a target profit level. In light of this point, the estimate in this section should be viewed as a lower bound on the actual farm price reduction required to lower the price of quotas to \$25,000 per unit.

Following the methodology proposed above, I find that in 2011 a quota price of \$25,000 per unit in Quebec would be supported by a farm price of \$66.00 per hectolitre. In 2011, Quebec dairy farmers received an average price of \$73.31 per hectolitre for farm milk,⁴⁴ and thus lowering the price to \$66.00 would represent a 9.97% reduction in the farm price. Figure 3.7 applies the same methodology and plots the target price of quotas as a function of the farm price of milk. The graph illustrates that any targeted quota price could be attained through adjusting the farm price of milk.

The impact of a 9.97% farm price reduction can be contextualized by considering the effect on Quebec dairy farmers' profit margin. Statistics Canada's Whole Farm Database is used to estimate the profit margin for Quebec dairy farmers, Canadian dairy farmers excluding Quebec dairy farmers, and all other animal product farmers excluding dairy farmers (this last category includes beef cattle, hog, poultry, and egg

⁴⁴Annual average calculated from FPLQ monthly reference price.



Figure 3.7: Target Dairy Quota Price as a Function of the Farm Price of Milk

farmers).⁴⁵ Figure 3.8 illustrates that dairy farming is the most profitable animal product industry in Canada. Over the period 1993-2011, the average profit margin of Canadian dairy farmers is over 11 percentage points higher than the average profit margin of other Canadian animal product farmers. Figure 3.8 also highlights that dairy farming in Quebec is more profitable than in the rest of Canada.

Holding constant the aggregate quota holding of each producer, a 9.97% reduction in the farm price results in a 9.97% reduction in farmers' revenue from dairy production. In 2011, a 9.97% reduction in revenue from dairy products and subsidies would have reduced Quebec dairy farmers' profit margin from 16.81% to 9.72%. Despite the magnitude of this decline, the profit margin of Quebec dairy farmers would remain 7.34 percentage points higher than the 2011 profit margin of other Canadian animal product farmers. However after the reduction, the profit margin of Quebec dairy

⁴⁵It is important to recognize that there are considerable differences in capital costs across the animal product industries and therefore a straight comparison of the ratio of net operating income to operating revenue would be misleading. To address this issue I define the profit margin as the ratio of net operating income adjusted for capital cost allowance to total operating revenue. Note that a more general definition of operating revenue and net operating income is used in this section of the paper as compared with the specification in equation (3.31). To facilitate comparison of the profit margin across different animal product industries total operating revenue and total operating costs are used (i.e. operating revenues and costs are inclusive of all farm enterprises).

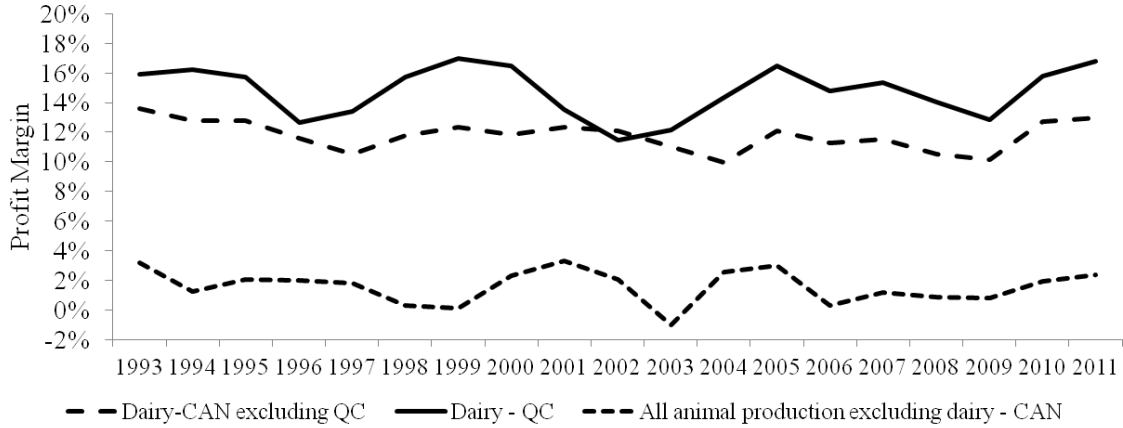


Figure 3.8: Profit Margins in Canadian Animal Product Farming

Source: CANSIM table 20048

farmers would be 3.28 percentage points lower than the 2011 average profit margin of dairy farmers in the rest of Canada.

The model developed in section 3 can be used to estimate the effects on the minimum viable scale of production that result from removing the price ceiling and lowering the farm price of milk. I begin by estimating the minimum viable scale in 2011 under price ceiling legislation by calibrating equation (3.27). To simulate the effect of removing the price ceiling, I impose market clearing for dairy quotas by evaluating equation (3.27) at the equilibrium price $P = \bar{\pi}/(r + \lambda)$ (this is equivalent to estimating the minimum viable scale by equation (3.26)). Finally, to estimate the effect of removing the price cap *and* lowering the farm price by 9.97%, I reduce $\bar{\pi}$ to account for the farm price reduction and re-evaluate the minimum scale using equation (3.26).

The minimum viable scale is estimated for 2011 using the following calibration: P is set equal to the 2011 ceiling price (converted to dollars per hectolitre), y is set to the Statistics Canada median household income for Quebec⁴⁶ in 2011; r is set to

⁴⁶Median household income data for Quebec is sourced from Statistics Canada CANSIM series

the prime interest rate plus two percent for 2011; the policy risk parameter λ is set equal to its econometrically estimated value for the period 1999-2005; η is set equal to $2/3$;⁴⁷ finally $\bar{\pi}$ is set to its 2011 value as defined by equation (3.31).⁴⁸ In 2011, the minimum scale of production is estimated to be 20.16 units of quotas, which roughly equates to a dairy herd of 20 cows. To simulate the effect of removing the price cap, I re-evaluate the minimum scale of production using equation (3.26) at the same calibration values and find that the minimum viable scale increases to 22.18 units of quotas. Therefore the price ceiling is estimated to have reduced the minimum viable scale of production by approximately 10%. Finally, the combined effect of removing the price ceiling and lowering the farm price by 9.97% is estimated by lowering the implicit farm price of milk, R/\tilde{Q} , by 9.97%, which by equation (3.31) results in a decrease in $\bar{\pi}$. Under this policy, the minimum scale of production is estimated to be 27.13 units of quotas, which represents a 34.6% increase as compared with the estimated minimum scale of production under price ceiling legislation in 2011.

Estimates of the minimum scale of production can also be used to interpret the pattern of net exit of Quebec dairy farms during the past two decades. Figure 3.9 plots the rate of net exit of dairy farms and the minimum viable scale of production during the period 1994-2011. The rate of net exit is calculated using the data on the population of Quebec farmers that is presented in Table 3.6. The minimum viable scale is estimated for the price ceiling period (2007-2011) using equation (3.27), and equation (3.26) for the period 1994-2006. The calibration values used in each

v41950213.

⁴⁷This parameterization reflects the precedent set by the United States' tobacco (2004) and peanut (2002) production quota buyout programs. Based on the evidence presented in Womach (2003), these programs paid farmers an amount ranging from approximately 60% to 74% of the asset value of their quotas.

⁴⁸The minimum viable scale is calculated in hectolitres per year and then converted to units of quotas (kilograms of butterfat per day) using a province specific conversion factor sourced from the Canadian Dairy Information Centre.

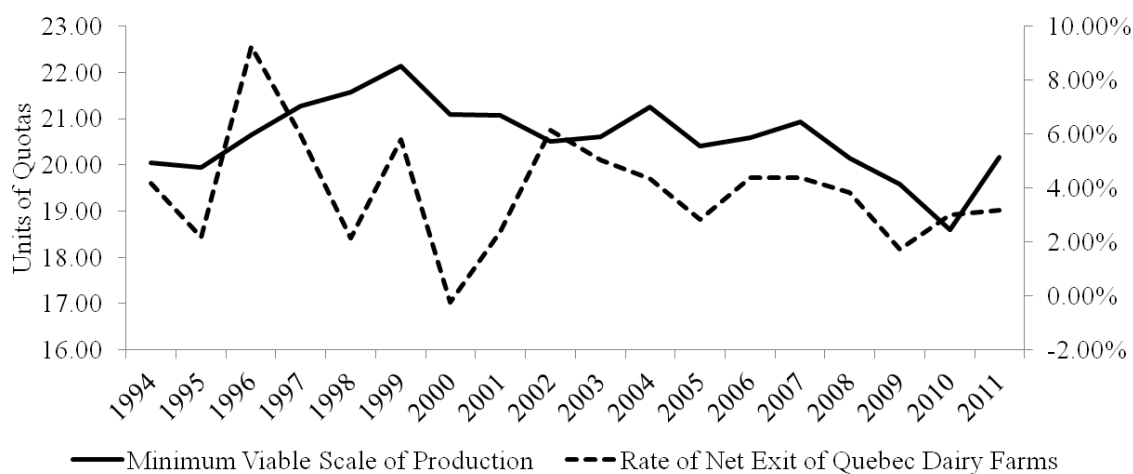


Figure 3.9: The Minimum Viable Scale of Production and Rate of Net Exit of Quebec Dairy Farms

respective year correspond with the specification described in the previous paragraph for the year 2011.⁴⁹ The average annual rate of net exit for Quebec dairy farms during the period 1994-2011 was 3.92%. As noted in my discussion of Table 3.6, the long run trend of net exit is also observed in the United States and is unlikely to be attributable to supply management. However, Figure 3.9 illustrates that the fluctuations around the trend of net exit appear to be positively correlated with the minimum viable scale, especially during the period 2003-2011. The correlation coefficients between the two time series for the periods 1994-2011 and 2003-2011 are 0.21 and 0.65 respectively. Changes in the minimum viable scale result from variation in factors that are exogenous to the sector (e.g. interest rate changes) and from policy changes that are determined within the supply management system. The positive correlation between the two series in Figure 3.9 suggests that farmers' exit decisions are responsive to supply management policy changes. For example, both the minimum viable scale of production and the net rate of exit declined following

⁴⁹The parameter λ is calibrated to the econometric estimates 0.1269, 0.0771, and 0.0463 for the periods 1994-1995, 1996-1998, and 1999-2011 respectively.

the introduction of the price ceiling in 2007. This suggests that the price ceiling has discouraged small farmers from exiting, thereby reducing the rate of decline in the number of dairy farmers in Quebec.⁵⁰

3.5.3 Policy Options for Reducing the Farm Price of Milk and Quota Prices

In this section I consider two policy experiments that could generate a farm price reduction of the magnitude required to reduce quota prices to \$25,000 per unit. The first policy considers reducing the farm price through increasing the TPQ. In this policy experiment I assume that the demand for raw milk exhibits constant own-price elasticity of demand. The functional form of the demand equation is $\tilde{Q}_t = a(P_t^m)^{-\epsilon}$, where $-\epsilon$ is the own-price elasticity of demand for raw milk. Define g_t^* as the growth in the TPQ that yields a targeted farm price of P_t^{m*} . Given ϵ and a reference price P_{t-1}^m it is possible to determine g_t^* from the demand equation:

$$P_t^{m*} = \left(\frac{a}{\tilde{Q}_t}\right)^{\frac{1}{\epsilon}} = \left(\frac{a}{Q_t(1+g_t^*)}\right)^{\frac{1}{\epsilon}} = \left(\frac{a}{Q_t}\right)^{\frac{1}{\epsilon}} \frac{1}{(1+g_t^*)^{\frac{1}{\epsilon}}} = P_{t-1}^m \frac{1}{(1+g_t^*)^{\frac{1}{\epsilon}}} \quad (3.33)$$

Where the final equality uses the definition of $\tilde{Q}_{t-1} = Q_{t-1}(1+g_{t-1}) = Q_t$. I specify $\epsilon = 0.72\epsilon_I + 0.28\epsilon_F$, where ϵ_I and ϵ_F are the own-price elasticities of industrial and fluid milk respectively. The weights of 0.72 and 0.28 are set equal to the fraction of Quebec's TPQ that comes from industrial and fluid milk respectively. The values of ϵ_I and ϵ_F are set to 0.5 and 0.1 following Meilke et al. (1998). Using this parameterization and the FPLQ's 2011 reference price of 73.31, I find that the farm price of milk could be lowered to \$66.00 through a 4.16% expansion in the TPQ.

⁵⁰I am grateful to an anonymous referee who recommended extending the model to address the issue of farmer exit.

The final counterfactual experiment in this section considers the effect of trade liberalization on the farm price of milk and the price of dairy quotas. Abbassi et al. (2008) simulate the impacts of various trade liberalization scenarios in the Canadian dairy industry using data from 2003-04. Abbassi et al. (2008) assume that the aggregate supply of raw milk remains constant in their moderate short-run trade liberalization scenario. This scenario involves a 25% reduction in bound tariffs, and a 50% increase in the minimum access commitment (MAC).⁵¹ The CETA *agreement-in-principle* provides the European Union with 16,000 tonnes of new access for cheese in the Canadian market, (Canada, 2013). This represents a 78% increase in the MAC for cheese, however CETA entails no additional European market access for other dairy products (milk, yogurt, butter, ice cream, etc.). The TPP may result in further increases in MAC levels, however as Gifford (2012) has commented, the TPP will likely require adjustment rather than elimination of supply management in Canada. Therefore, the trade liberalization scenario considered by Abbassi et al. (2008) appears to be comparable to a moderate outcome to the prospective TPP and CETA agreements.⁵² Abbassi et al.'s (2008) simulation of the above scenario results in a 10.3% reduction in the farm price of milk in the province of Quebec. Under this same scenario Abbassi et al. (2008) estimate that the welfare of Canadian dairy farmers is reduced by 25.2%, while total Canadian welfare increases by 0.4%.

Drawing on Abbassi et al.'s (2008) results and using the methodology presented above in equation (3.32), I estimate that a 10.3% reduction in the farm price of milk would result in a farm price of \$65.76 and reduce the 2011 price of dairy quotas in

⁵¹Trade of Canadian dairy products is regulated by tariff rate quotas (TRQs). TRQs involve a MAC and two tariff rates, an in-quota and over-quota tariff. The in-quota tariff is applied to imports not exceeding the MAC, while the over-quota tariff is applied to imports exceeding the MAC.

⁵²The 25% reduction in bound tariffs considered by Abbassi et al. (2008) appears to be less likely in the context of the TPP and CETA agreements. However, the authors note that the effect of the over-quota tariff decrease is negligible in this trade liberalization scenario, as the magnitude of the decrease is not large enough to trigger imports above the MAC.

Quebec to \$24,820 per unit. Thus, it is likely that the price of dairy quotas in Quebec would drop to a level roughly equivalent to the current ceiling price following a 25% reduction in bound tariffs and 50% increase in the MAC, assuming that the aggregate supply of raw milk in Quebec is held constant following trade liberalization.

3.6 Conclusions

The price of dairy quotas has risen dramatically over the past two decades. The empirical results in this paper provide evidence in support of the theory that reduced policy risk and improved borrowing conditions in the 1990s put upward pressure on the price of dairy quotas. However, the results also indicate that quota prices may have risen due to increases in the average profit per unit of quota over the past twenty years.

I estimate that dairy quotas in Quebec in 2011 would have traded at a price of \$29,923 per unit in the absence of the price ceiling. My results indicate that lowering the market valuation of quotas to \$25,000 per unit would have required a 9.97% reduction in the farm price of milk. In 2011, a 9.97% reduction in operating revenue would have reduced Quebec dairy farmers' profit margin to 9.72%. While this is the lowest level in recent history, the margin would remain 7.34 percentage points higher than the 2011 profit margin of other Canadian animal product farmers.

The recent implementation of quota price ceilings in Canada's two largest dairy producing provinces have introduced a new source of inefficiency in the Canadian dairy sector. From the perspective of economic theory, lowering the farm price of milk is preferred to a price cap as a method for lowering the price of quotas. Consumers also stand to benefit from a reduction in the farm price of milk, as the profitability of dairy farming under supply management is reflected in higher retail prices for Canadian

dairy products. While I have not estimated the welfare effects of a 9.97% decrease in the farm price of milk, the inelasticity of consumer demand for dairy products suggests that there would be large gains in consumer surplus from the counterfactual policy experiments considered in this paper. The counterfactual experiments in section 5 indicate that the price of quotas could be reduced to the ceiling price level through a 4.16% expansion of the TPQ, or through a trade liberalization scenario involving a 25% reduction in bound tariffs and 50% increase in the MAC. The farm price reduction resulting from these policies would improve the competitive position of the Canadian dairy sector, which is imperative given that recent research suggests that the sector is not prepared to compete in international markets, (Kimura and Thi, 2013). The results of this paper highlight the complementarity between policy reform and improved efficiency in the Canadian dairy sector.

Chapter 4: Trade Liberalization, Selection, and Productivity in a Supply-Managed Economy

4.1 Introduction

This paper focuses on the productivity effects of the Commercial Export Milk Program (CEM) in the Canadian dairy sector. This program was active from 2000-2003 and allowed farmers to sell milk destined for export directly to Canadian processors. Recently, the Canada-EU Comprehensive Economic and Trade Agreement (CETA) and the Trans-Pacific Partnership (TPP) have resulted in modest liberalization of Canada's supply-managed agricultural industries. These small concessions may nevertheless signal the path for more substantial reforms in the future. It is difficult to predict the impacts of trade liberalization in these industries since farmers' responses to competitive price movements are unobserved under supply management. Under the CEM program the farm price of milk was approximately half of the administered supply-managed price. Therefore, studying the CEM program provides insight into the potential effects of trade liberalization and competitive pricing in Canada's supply-managed agricultural industries.

In 1999 the WTO's Dispute Settlement Body ruled that Canada must reform its dairy export mechanism. Canada responded by introducing the CEM program and a number of other new export policies for the Canadian dairy industry. The CEM program permitted farmers to sell milk directly to Canadian processors provided that the final goods were destined for export. Importantly, farmers did not require production quotas for milk sold through the CEM program. Participation in the program was voluntary, and the farm price of milk was negotiated directly between the farmer and the processor. The CEM program created a dual marketing stream for raw milk that circumvented the highly regulated system of marketing under supply management. Therefore, farmers who participated in the program effectively faced a competitive marketing environment. The CEM program was canceled in 2003 after a 2002 WTO ruling concluded that the CEM program constituted an export subsidy for Canadian dairy processors.

The data for this paper comes from the Canadian Census of Agriculture (1996, 2001, 2006) and the Quebec Federation of Management Clubs' Agritel database (1993-2006). I begin by estimating farm-level total factor productivity (TFP) using a range of firm-level production function estimation methodologies. I then estimate the 'export premia' regressions that are routinely reported in the empirical trade literature following the methodology of Bernard and Jensen (1999). These regressions show a positive correlation between farm-level TFP and participation in the CEM program. This positive correlation persists when the sample is restricted to observations in the pre-policy period, which provides evidence of the self-selection of highly productive farms into the CEM program. To test if participation in the CEM program affected farm-level TFP, I use difference-in-difference methodology and inverse propensity weighting following the approach of Hirano et al. (2003). I find no evidence of a statistically significant causal effect of CEM participation on farm-level TFP.

This paper contributes to the empirical trade literature on firm heterogeneity and export participation. While there is a large literature documenting the correlation between export participation and firm-level productivity, relatively few papers have tested for causality in this relationship. Examples include: Clerides et al. (1998), Bernard and Jensen (1999), Van Biesebroeck (2005), and De Loecker (2007). The majority of papers in this literature focus on a single sector: manufacturing. To my knowledge, this is the first paper to test for causation between trade participation and productivity using farm-level data in the agriculture sector. Understanding this relationship is important as tariff rates in agriculture are notoriously high in comparison with other industrial goods. By focusing on the Canadian dairy industry, this paper provides insight into the gains from trade in the agricultural sector where a great deal of trade liberalization has yet to occur.

Earlier papers examining the CEM program include Doyon et al. (2006) and Turvey et al. (2003). However, these papers focus primarily on domestic production quota pricing, and neither use farm-level data or consider the effect of CEM participation on farm-level productivity. More recently, Barichello et al. (2014) argues that the Canadian dairy industry stands to benefit from trade liberalization by showing that the potential gains from export opportunities in emerging markets outweigh the losses associated with eliminating supply management. However Barichello et al.'s (2014) analysis does not account for the added welfare gains from trade liberalization that arise from farm-level heterogeneity. The theoretical framework of Melitz (2003) shows that trade liberalization results in a re-allocation of output to the highly-productive firms that are competitive in international markets. This paper finds a positive correlation between farm-level TFP and participation in the CEM program. This result is consistent with Melitz (2003), as highly productive farmers self-selected into the CEM program when facing a CEM price that was much lower than the supply-managed

price. This finding also suggests that full trade liberalization of the Canadian dairy industry would result in welfare gains from the re-allocation of output to the most productive farmers who would remain competitive at a lower farm price of milk.

The remainder of this paper is organized as follows: Section 4.2 describes the data, Section 4.3 reports the farm-level TFP estimates, Section 4.4 presents the research design and main results of the paper, and Section 4.5 concludes.

4.2 Data

The main source of data for this paper is the Agritel panel database, which is generated from a farm-level survey completed annually by the Federation of Management Clubs in the province of Quebec. In Section 4.3, I describe the variables from the Agritel database that are used to estimate the farm-level production function and TFP. In this section, I focus on describing the variables that relate to the CEM program. My previous research indicates that the Agritel database is not representative of the population of dairy farms in Quebec. Over the period 1993-2011, the average revenue of farms in the Agritel database was 42 percent higher than for the population of dairy farms in Quebec, (Chernoff, 2015). To correct for the unrepresentative nature of the sample, I construct sample weights using farm-level data from the Canadian Census of Agriculture for the years 1996, 2001, 2006.⁵³

Table 4.8 reports summary statistics on farmers' participation in the CEM pro-

⁵³The weights are constructed using the following procedure: In each Census year a tabulation of frequencies of the number of cows per farm is calculated from the Census farm-level data. That is, I count the number of farms with herd sizes within discrete intervals (the classes are: 1-10 cows per farm, 11-20 cows per farm, etc). I then calculate the average of the frequencies in each discrete interval over the three Census years (1996, 2001, 2006). Next I make a corresponding tabulation of frequencies for each year over the period 1996-2006 using the Agritel data. I then calculate the average of the frequencies in the Agritel data over the period 1996-2006. For each interval, the sample weight is calculated to equate the frequency counts from the Agritel tabulation with the Census tabulation.

gram between 2000 and 2003. The second column in Table 4.8 shows the total number of farmers in the Agritel Survey during each of the four years of the program. The third column shows the weighted percentage⁵⁴ of these farmers who participated in the CEM program, and thus provides an estimate of the percentage of Quebec farmers participating in the program. In the years 2001 and 2002, participation in the program peaked at 38.3 and 32.5 percent respectively. Whereas few papers have documented farm-level participation in export markets for agricultural industries, it is useful to compare the statistics in Table 4.8 with firm-level export data for the manufacturing sector during the same era.⁵⁵ Bernard et al. (2007) find that 18 percent of U.S. manufacturing firms participated in export markets in 2002. By this comparison, farm-level participation in the CEM program at the extensive margin was relatively high.

Column four indicates that milk marketed through the CEM program accounted for only a small percentage of output (output is measured in hectolitres (hl) of milk). At the height of the program in 2001, the average output marketed through the CEM program was only 2.57 percent. Bernard et al. (2007) estimate that mean exports accounted for approximately 14 percent of output in the U.S. manufacturing sector in 2002. Therefore farm-level participation in the CEM program at the intensive margin of trade was relatively low.

Columns five and six of Table 4.8 compare the CEM price, P_{CEM} , with the price farmers receive under supply management, P_{SM} . A comparison of these prices reveals that the average CEM price was 54 percent lower than the supply-managed price

⁵⁴Henceforth all averages are weighted by the Census population weights unless otherwise noted.

⁵⁵When drawing comparisons to the empirical trade literature, it is important to note how participation in the CEM program differs from conventional notions of exporting. In order to participate in the CEM program, the milk procured by processors under a CEM contract needed to be destined for export markets. Therefore farmers only *indirectly* participated in export markets through processors and the CEM contracts. This differs from most studies in the empirical trade literature that study firm's that directly participate in export markets.

during the four years of the program. While farmers received a much lower price for their milk under the CEM marketing scheme, they avoided the costs associated with purchasing production quotas. Production quotas account for a large fraction of farmers' assets, and the high cost of quotas can present a barrier to output growth.

Table 4.8: Quebec Farm Participation in the Commercial Export Milk Program

(1) Year	(2) Sample	(3) CEM Participants	(4) <u>Exports</u> Output	(5) P_{CEM}	(6) P_{SM}	(7) CEM Fees	(8) SM Fees
2000	1,315	9.54%	0.38%	\$27.40	\$56.55	\$1.25	\$3.17
2001	1,287	38.30%	2.57%	\$29.53	\$57.43	\$2.17	\$3.49
2002	1,217	32.50%	2.29%	\$25.32	\$58.53	\$2.13	\$3.77
2003	1,169	7.50%	0.28%	\$24.16	\$60.34	\$1.76	\$3.88

Column two is the Agritel sample size; column three is the percentage of farmers in the sample that participated in the CEM program; column four is the average percentage of farmers milk marketed through the CEM; column five is the average price farmers received for milk sold through the CEM program; column six is the monthly reference price sourced from the Fédération des Producteurs de Lait du Québec; and columns seven and eight are the average milk marketing fees associated with the CEM program and supply management respectively. Dollar values are in Canadian dollars per hectolitre (hl). Columns three to five, seven, and eight are weighted sample averages using the Agritel database and the sample weights constructed from the Canadian Census of Agriculture.

The high percentage of farmers that marketed milk through the CEM program can be explained by the ease at which farmers could participate in the program. In order to participate, farmers simply needed to respond to the CEM contracts posted by Canadian dairy processors. The posted contracts listed the offer price and quantity demanded. Turvey et al. (2003) note that processors adopted a practice of posting contracted quantities far in excess of expected supply as a means of signaling their intention to accept all milk supplied at the posted price. For example, in November of 2000 Nestlé Canada posted a request for 20 million litres per day at \$33.4658 per hectolitre. Turvey et al. (2003) note that only 1.51 % of the total quantity demanded by processors under CEM contracts in Ontario were filled.⁵⁶ Participation in the CEM

⁵⁶The monthly CEM program data used by Turvey et al. (2003), spanning July 2000 to November 2001, does not cover the complete duration of the program.

program was also facilitated by lower marketing costs as compared with the supply management. Column seven presents the average marketing fees paid by farmers participating in the CEM program, and column eight presents the average marketing fees associated with marketing milk through supply management. Marketing fees under the CEM program were 49 percent below the marketing fees farmers paid to market their milk through supply management. It is worth emphasizing that the supply management marketing fees listed in column eight must be paid by farmers in addition to the costs associated with acquiring production quotas.

4.3 Total Factor Productivity Estimation

In this section I describe the production technology, TFP estimation methodology, and results. I assume that farmers produce milk using a Cobb-Douglas production function. The natural logarithmic representation of the production technology is given by the following equation:

$$Output_{i,t} = \beta_0 + \beta_1 Labor_{i,t} + \beta_2 Concentrates_{i,t} + \beta_3 Forages_{i,t} + \beta_4 Cows_{i,t} \quad (4.34)$$

$$+ \beta_5 Capital_{i,t} + u_{i,t}, \quad u_{i,t} = \Omega_{i,t} + \eta_{i,t}$$

Output and all inputs in equation (4.34) are expressed in logarithmic form. The variable $Output_{i,t}$ is the hectolitres of milk produced by farmer i in year t . The production function includes labor input, which is measured in physical units using the standardized method employed in the Agritel survey. The production function variables also include physical units of concentrates and forages. These variables represent the two largest variable inputs in dairy production as measured by their share in the average variable cost of production during the observation period, 1993-

2006. The next variable included in the production function is the herd size, which is measured by the number of dairy cows used by the farmer. The final variable in the production function is physical capital, which is measured by the real value of total farm assets net of dairy production quotas. To convert nominal to real values I use a capital price index constructed from CANSIM Table 383-0032 that is specific to NAICS Industry 11, Agriculture, Forestry, Fishing and Hunting.

I estimate equation (4.34) by ordinary least squares, fixed effects, and by a variation of the methodology of Olley and Pakes (1996). Following Olley and Pakes (1996) I assume the error term, $u_{i,t}$, is composed of two additive components, $\Omega_{i,t}$ and $\eta_{i,t}$. The first term, $\Omega_{i,t}$ is a productivity shock that is observed by the farmer but not the econometrician. The second term, $\eta_{i,t}$, is a second productivity shock that is unobserved by both the farmer and econometrician. After observing $\Omega_{i,t}$ the farmer is assumed to simultaneously choose the level of output and variable inputs to maximize profits. The simultaneity of this decision results in an upward bias on the variable input coefficient estimates. If the productivity shock $\Omega_{i,t}$ is assumed constant over time, then fixed effects estimation can be used to account for unobserved farm-level heterogeneity. Under the more realistic assumption that farmers' productivity varies over time, Olley and Pakes (1996) estimation can be used to account for the simultaneity in input and output decisions.

For the purposes of Olley and Pakes (1996) estimation, $Labor_{i,t}$, $Concentrates_{i,t}$, $Forages_{i,t}$, and $Cows_{i,t}$ are specified as variable inputs, while $Capital_{i,t}$ is specified as fixed in any given year. I follow Olley and Pakes (1996) in using investment as a proxy for productivity. The investment variable is specified in logarithmic form and therefore all farmers with non-positive investment are not used in the Olley and Pakes (1996) estimation procedure. This can result in sample selection bias if a large number of farmers have non-positive investment. In my sample 94.76 percent of observations

had positive levels of investment, and therefore the effect of sample selection bias is expected to be small.

For the Olley and Pakes (1996) estimator, I follow the one-step estimation procedure proposed by Wooldridge (2009). This procedure has the advantage of easily accommodating the sample weights that were described in Section 4.2. A second advantage of the Wooldridge (2009) estimator is that it addresses the critique raised by Akerberg et al. (2006). Akerberg et al. (2006) note that the Olley and Pakes (1996) estimator suffers from collinearity problems in the econometric model. Wooldridge's (2009) one-step estimator can be used to implement Olley and Pakes's (1996) estimator in a manner that addresses the critique of Akerberg et al. (2006). A limitation of the Wooldridge (2009) estimator is that it does not account for the selection bias associated with firm-exit that is explicitly modeled in Olley and Pakes (1996). However, Olley and Pakes (1996) find that there are only minor efficiency gains from controlling for exit when an unbalanced panel is used in estimation. A survey by Van Beveren (2010) of the recent firm-level productivity estimation methodologies reiterates the finding that correcting for firm survival has little effect on the estimated coefficients of the production function. Furthermore, I do not have a reliable measure of farmer exit as it is impossible to distinguish farmer exit from survey attrition in the Agritel panel database.

The Wooldridge (2009) estimator entails using a control function to account for unobserved farm-level productivity. In particular, I augment equation (4.34) with the following third order polynomial control function in capital and investment:

$$g(\text{Capital}_{i,t}, \text{Invest}_{i,t}, t) = \gamma_1 \text{Capital}_{i,t-1} + \gamma_2 \text{Capital}_{i,t-1}^2 + \gamma_3 \text{Capital}_{i,t-1}^3 \quad (4.35)$$

$$+ \gamma_4 \text{Invest}_{i,t-1} + \gamma_5 \text{Invest}_{i,t-1}^2 + \gamma_6 \text{Invest}_{i,t-1}^3 + \gamma_7 \text{Invest}_{i,t-1} \text{Capital}_{i,t-1}$$

$$+\gamma_8 Invest_{i,t-1}^2 Capital_{i,t-1} + \gamma_9 Invest_{i,t-1} Capital_{i,t-1}^2$$

I follow Petrin et al. (2013) and Wooldridge (2009) in specifying the control function as a third order polynomial function. This higher order polynomial specification of the control function is used to accommodate a flexible relationship between unobserved productivity and lagged capital and investment. The capital and investment variables in equation (4.35) are expressed in logarithmic form. The Wooldridge (2009) estimator also requires specifying instrumental variables for the variable inputs. I follow Wooldridge (2009) and Petrin et al. (2013) by using the first lag of the labor, concentrates, forages, and cows as instrumental variables.

Table 4.9: Production Function Estimation Results

Variable	OLS	Fixed Effects	Olley & Pakes
Labor	0.00104 (0.0156)	0.0214** (0.0107)	0.00867 (0.00905)
Concentrates	0.347*** (0.0134)	0.223*** (0.0119)	0.440*** (0.0249)
Forages	0.0864*** (0.0139)	0.0453*** (0.00901)	0.133*** (0.0344)
Cows	0.574*** (0.0307)	0.774*** (0.0608)	0.403*** (0.0405)
Capital	0.0554*** (0.0107)	0.0961*** (0.0214)	-0.0129 (0.0263)
Observations	17,832	17,832	13,556
R-squared	0.929	0.790	0.922
Number of unique farmers		2,548	

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable and all explanatory variables are expressed in natural logarithm form. Estimates use sample weights constructed from the Canadian Census of Agriculture.

Table 4.9 reports the results of estimating equation (4.34) by OLS, fixed effects (henceforth FE), and by Olley and Pakes's (1996) estimation using the one-step methodology of Wooldridge (2009) (henceforth OP). The coefficient estimates in Ta-

ble 4.9 are used to construct farm-level TFP as defined by the regression residual from equation (4.34). In general, the magnitude and significance of the coefficient estimates vary across the three estimation methodologies. In the remainder of the paper, separate regressions will be run using the OLS, FE and OP TFP estimates to demonstrate the robustness of the results to the different production function estimation methodologies.

4.4 CEM Participation and TFP

4.4.1 Export Premia Regressions

In this section I present the export premia regressions that are commonly reported in the empirical trade literature following the methodology of Bernard and Jensen (1999). The export premia regression takes the following form:

$$TFP_{i,t} = \beta_0 + \beta_1 xstatus_i + Z_{i,t}\delta + u_{i,t} \quad (4.36)$$

The dependent variable, $TFP_{i,t}$, is the natural logarithm of TFP. A separate regression is run for each of the three different methodologies used to estimate TFP in Section 4.3: OLS, FE, and OP. The vector $Z_{i,t}$ contains variables that control for year fixed effects and the natural logarithm of output to control for economies of scale. The variable of interest is the dummy variable indicating farmers' participation status in the CEM program. This variable takes a value of one if farmer i participated in the CEM program at least once during the period 2000-2003. The coefficient on the CEM participation variable, β_1 , is the key coefficient of interest. This coefficient is referred to as the export premia and measures the correlation between CEM participation and TFP, holding constant the other control variables.

Table 4.10: Export Premia Regressions - 1993-2006

Dependent Variable	TFP_{OLS}	TFP_{FE}	TFP_{OP}
<i>xstatus</i>	0.0137*** (0.00309)	0.0127*** (0.00349)	0.0119*** (0.00279)
Output	0.0608*** (0.0175)	0.00958 (0.0200)	0.125*** (0.0145)
Observations	17,832	17,832	17,832
R-squared	0.099	0.045	0.251

Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Each regression model includes dummy variables for year fixed effects. Estimates use sample weights constructed from the Canadian Census of Agriculture. The CEM participation variable, *xstatus* is an indicator variable that takes a value of one if farmers participated in the CEM program at least once during the period 2000-2003.

The results of the export premia regressions for the observation period 1993-2006 are reported in Table 4.10. In all three specifications the regressions show a positive relationship between CEM participation and TFP, as the export premia coefficient is statistically significant at the 1 percent level. The results also show a positive and highly significant relationship between farm size and TFP in the OLS and OP specifications. This is consistent with prior research indicating that there are increasing returns to scale in dairy farming.⁵⁷

The theoretical trade literature has offered a number of explanations for the positive correlation between trade participation and productivity. In Melitz (2003) firms must pay a fixed cost to enter export markets which leads to selection of the most productive firms into trade. To examine if selection partially explains the positive correlation between CEM participation and productivity, I re-estimate the export premia regressions for the period *before* the CEM program was implemented, 1993-1999. The results are reported in Table 4.11. Once again, in all three regressions the

⁵⁷For example, Mosheim and Lovell (2009) find evidence of increasing returns to scale in US dairy farming. Similarly, Moschini (1988) finds that there are increasing returns to scale in dairy farming in Ontario.

export premia coefficient is positive and statistically significant at the 1 percent level. This provides strong evidence that selection at least partially explains the positive correlation between CEM participation and farm-level productivity.

It is useful to benchmark the estimated export premia for the Canadian dairy farming industry to that of the U.S. manufacturing sector during the same era. Bernard et al. (2007) report that the TFP export premia for the U.S. manufacturing sector ranged from 0.02 to 0.05 which is higher than the export premia reported for the Canadian dairy sector in Table 4.10. The summary statistics presented in Table 4.8 may help to explain this result. In section 4.2 I noted that the marketing fees for the CEM program were low relative to the marketing fees incurred from selling milk domestically through supply management. Interpreted in the context of the Melitz (2003) model, the magnitude of the selection effect may have been relatively small for the CEM program since the costs of participating in the program was low.

Table 4.11: Export Premia Regressions - 1993-1999

Dependent Variable	TFP_{OLS}	TFP_{FE}	TFP_{OP}
<i>xstatus</i>	0.0135*** (0.00324)	0.0125*** (0.00346)	0.0125*** (0.00323)
Output	0.0992*** (0.0248)	0.0570** (0.0272)	0.154*** (0.0218)
Observations	9,769	9,769	9,769
R-squared	0.133	0.063	0.256

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Each regression model includes dummy variables for year fixed effects. Estimates use sample weights constructed from the Canadian Census of Agriculture. The CEM participation variable, *xstatus* is an indicator variable that takes a value of one if farmers participated in the CEM program at least once during the period 2000-2003.

4.4.2 Identification

In this section I turn to the question of whether or not participation in the CEM program affected farm-level productivity. A naive approach to addressing this question would be to use a standard difference-in-difference research design. This would involve supplementing the export premia regression model with interaction terms as follows:

$$\begin{aligned} TFP_{i,t} = & \beta_0 + \beta_1 xstatus_i + \beta_2 During_t + \beta_3 xstatus_i x During_t \\ & + \beta_4 After_t + \beta_5 xstatus_i x After_t + Z_{i,t}\delta + u_{i,t}, \end{aligned} \quad (4.37)$$

where *During* and *After* are indicator variables that take a value of one if the period is 1993-1999, 2000-2003, and 2004-2006 respectively.

The problem with a standard difference-in-difference research design is that the treatment (participation in the CEM program) is not randomly assigned. The results in Table 4.11 indicate that there were ex-ante differences between the farmers who did and did not participate in the CEM program. In particular, the results in the previous section indicate that the control group (farmers who did not participate in the CEM program) had lower TFP on average. Inverse propensity weighting can be used to improve the research design by increasing the ex-ante similarity between the treatment and control group. After the inverse propensity weights have been applied, the weighted average observable characteristics of the two groups are nearly identical for the period before the CEM program. Using the difference-in-difference regressions in conjunction with the inverse propensity weights identifies the coefficient of interest, β_3 , that measures the effect on TFP from farm participation in the CEM program.

To calculate the inverse propensity weights I follow the procedure introduced by

Hirano et al. (2003). First I estimate the probability $\hat{P}_{i,t}$, which is the probability that farmer i in period t belonged to the treatment group conditional on the farmer's observable characteristics. The observable characteristics include the farmer's TFP, herd size, and age.⁵⁸ The probability $\hat{P}_{i,t}$ is estimated from the following probit regression, which is estimated for the period *before* the CEM program was implemented (1993-1999):

$$xstatus_i = \gamma_0 + \gamma_1 TFP_{i,t} + \gamma_2 Cows_{i,t} + \gamma_3 age_{i,t} + u_{i,t} \quad (4.38)$$

Next I calculate the inverse propensity weights using the following equation:

$$w_{it} = \frac{xstatus_i}{\hat{P}_{i,t}} + \frac{1 - xstatus_i}{1 - \hat{P}_{i,t}} \quad (4.39)$$

Finally I create a new set of regression weights by multiplying the inverse propensity weights in equation (4.39) by the sample weights created from the Census of Agriculture. These new weights are used in estimating the difference-in-difference regression equation (4.37).

4.4.3 Difference-in-Difference Results

Table 4.12 presents the results from estimating equation (4.37) by OLS with the inverse propensity weights that were defined in the previous section. The regressions are run with the OLS, FE, and OP TFP estimates from Section 4.3.

The key parameter of interests is the coefficient for the interaction variable, $xstatus_i \times During_t$. This parameter captures the effect of the CEM participation on farm-level productivity. This coefficient is not statistically significant in each of

⁵⁸As in Section 4.3, TFP and herd size are expressed in logarithmic form.

Table 4.12: CEM Participation and TFP Regression Results

Dependent Variable	TFP_{OLS}	TFP_{FE}	TFP_{OP}
xstatus	0.0215*** (0.00337)	0.0209*** (0.00368)	0.0206*** (0.00325)
During	0.0357*** (0.00902)	0.0444*** (0.0101)	0.0454*** (0.00855)
After	0.0397*** (0.0129)	0.0460*** (0.0150)	0.0524*** (0.0115)
xstatus x During	-0.00156 (0.00505)	-0.00328 (0.00555)	-0.00147 (0.00486)
xstatus x After	-0.0121 (0.0105)	-0.0110 (0.0126)	-0.0145* (0.00852)
Output	0.0519*** (0.0181)	0.000593 (0.0208)	0.116*** (0.0149)
Observations	17,201	17,201	17,201
R-squared	0.092	0.046	0.247

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Estimates use inverse propensity weights combined with sample weights constructed from the Canadian Census of Agriculture. Each regression model includes dummy variables for year fixed effects. The CEM participation variable, *xstatus* is an indicator variable that takes a value of one if farmers participated in the CEM program at least once during the period 2000-2003.

the three specifications. This result indicates that participating in the CEM program did not have a statistically significant effect on farm-level TFP. This result may partially be explained by the fact that the total quantity of milk marketed through the CEM program was relatively small. As noted in Table 4.8, although a large fraction of farmers participated in the program, the total percentage of milk sold through the CEM program peaked in 2001 at only 2.57%.

The coefficient on the $xstatus_i$ variable is positive and statistically significant at the one percent level in each specification. This is consistent with the results of Section 4.4.1, and provides further evidence of positive-selection into the CEM program. Finally the coefficient on the interaction term, $xstatus_i \times After_t$, is not statistically significant under OLS or FE, and is negative and statistically significant at the ten percent level under the OP specification of TFP. This provides weak evidence that participants in the CEM program experienced a decline in their TFP after the program was canceled. A possible explanation for this result is that farmers may have participated in the CEM program as an alternative to increasing their dairy production quotas. When the CEM program was canceled, these same farmers may have had a drop in their TFP due to having production capacity in excess of their quota allocation.

4.5 Conclusion

Carter and Mérel (2014) and Barichello et al. (2014) have argued that farmers in Canadian supply-managed industries could potentially benefit from trade liberalization. Under full trade liberalization Barichello et al. (2014) predict that the Canadian dairy industry could achieve a cumulative growth rate of 5.8 percent annually. However, in Barichello et al.'s (2014) analysis the gains in producer surplus are limited to

those associated with scale economies and market expansion.

The findings of this paper suggest there may be additional pro-competitive gains from liberalizing the Canadian dairy industry arising from farm-level heterogeneity. Although the results show no evidence of a causal effect of the CEM program on productivity, the results indicate that the most productive farmers self-selected into the program. This finding is consistent with the theoretical predictions of the Melitz (2003) model of heterogeneous firms in trade. In the Melitz (2003) model the most productive firms self-select into export markets when prices fall in response to trade liberalization. Under the CEM program, farmers received a price that was approximately half of the domestic supply-managed farm price. However, the magnitude of the aggregate re-allocation of output to the highly-productive farmers that participated in the CEM program is likely to have been small. In fact, the total milk sold through the CEM program is estimated to have accounted for less than three percent of aggregate annual output in Quebec.

The small percentage of output sold through the CEM program can be explained by the fact that the program was voluntary and existed alongside the supply management marketing stream. The research of Abbassi et al. (2008) and Barichello et al. (2014) suggests that liberalizing the Canadian dairy industry would result in a significant drop in the farm price of milk. Unlike the CEM program, all farmers would face a lower world price under free trade. The results of this paper suggest that under trade liberalization, there would be welfare gains arising from the re-allocation of output to the most productive farmers who would remain competitive in international markets.

Chapter 5: Conclusion

Models of firm-heterogeneity are increasingly used in many fields of economics to study the relationship between firm-level and aggregate outcomes. Empirical and theoretical models have been developed to analyze how economic policy affects aggregate welfare and the distribution of output across firms with varying levels of productivity. The objective of this thesis is to use theoretical models of firm-heterogeneity to complete empirical analyses in economic history and agricultural economics.

In Chapter 2, I develop a framework to analyze the productivity and welfare gains from steam power adoption in the late nineteenth century. In the empirical application of the paper I use Canadian firm-level data and an instrumental variables approach to estimate the effect of steam-power on firm-level productivity. My results indicate that the use of steam power resulted in a 15.1 percent increase in firm-level productivity and a 3.0-5.2 percent increase in welfare. The empirical framework developed in this paper draws on methodology that has been used to estimate the welfare gains from international trade in models with heterogeneous firms. Comparing my results to this literature, I show that the estimated welfare gains from steam power adoption are approximately one and a half times as large as contemporary estimates of the welfare gains from trade.

In Chapter 3, I consider various policy alternatives to price ceiling legislation in the market for production quotas in the dairy farming sector in Quebec. I develop a

dynamic model of the demand for quotas with farmers that are heterogeneous in their marginal cost of milk production. The econometric analysis estimates a structural parameter of the theoretical model that is required for the counterfactual experiments. Using my econometric results and the modeled equilibrium price, I estimate the price of dairy quotas over the period 1993-2011. The counterfactual experiments indicate that the price of quotas could be reduced to the ceiling price through a 4.16 percent expansion of the aggregate supply of quotas, or through moderate trade liberalization of Canadian dairy products.

In Chapter 4, I examine the relationship between productivity and participation in the CEM program. I find a positive correlation between participation in the Commercial Export Milk Program (CEM) program and farm-level Total Factor Productivity (TFP). I then use a difference-in-difference research design with inverse propensity weights to test for causality in the relationship between participation in the CEM program and TFP. I find no statistically significant evidence that the CEM program affected farm-level TFP. A number of economists have argued that the Canadian dairy industry could benefit from trade liberalization through export market growth and returns to scale in production. My results suggest an additional channel through which trade liberalization would increase farm-level productivity. Under the CEM program farmers received a farm price of milk that was approximately half of the domestic price under supply management. My results indicate that highly productive farmers self-selected into participating in the program. This finding suggests that under trade liberalization there would be productivity gains from the re-allocation of output to the most productive farmers that would remain competitive under a lower farm price of milk.

Chapters 2-4 use models of firm-heterogeneity to complete empirical analyses in economic history and agricultural economics. While these models have been widely

used in various fields of economics, applications in economic history and agricultural economics have been less common. My findings demonstrate how incorporating firm heterogeneity provides new insights on the research questions considered in this thesis.

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Appendix A:

A.1. Proofs and Derivations from Chapter 2

Proof of Proposition 1

It is convenient to invert the demand curve and write the revenue of a firm operating with hand power as a function of demand:

$$r(\varphi, 1) = P^{\frac{\sigma-1}{\sigma}} q(\varphi, 1)^{\frac{\sigma-1}{\sigma}} \quad (5.40)$$

The above expression also defines firm-level revenue in the economy without steam power. Note that the revenue of a firm operating with productivity equal to the operating cut-off is equal in the economy with and without steam, $r(\varphi_1, 1) = r(\varphi'_1) = \sigma f$. Using this equality to evaluate equation (5.40) at φ_1 and φ'_1 yields:

$$\mathbb{W}/\mathbb{W}' = P'/P = q(\varphi_1, 1)/q(\varphi'_1, 1) \quad (5.41)$$

Finally, I use the fact that $r(\varphi_1, 1) = r(\varphi'_1) = \sigma f$ along with the definition of CES demand implies $q(\varphi_1, 1) = (\sigma - 1)f\varphi_1$ and $q(\varphi'_1, 1) = (\sigma - 1)f\varphi'_1$. Combining these expressions with equation (5.41) and substituting in the equilibrium values of the operating cut-offs, φ_1 and φ'_1 , completes the proof. \square

Derivation of the Distribution of Value Added for Firms with Steam Power

Define $F(r)$ as the CDF of value added for the subset of firms using steam power. By definition:

$$\begin{aligned} F(r) &= Pr(r(\varphi, \gamma_3) \leq r \mid r(\varphi_3, \gamma_3) \leq r(\varphi, \gamma_3)) = \frac{Pr(r(\varphi_3, \gamma_3) \leq r(\varphi, \gamma_3) \leq r)}{Pr(r(\varphi, \gamma_3) \geq r(\varphi_3, \gamma_3))}, \\ &= \frac{Pr(r(\varphi, \gamma_3) \leq r) - Pr(r(\varphi, \gamma_3) \leq r(\varphi_3, \gamma_3))}{Pr(r(\varphi, \gamma_3) \geq r(\varphi_3, \gamma_3))} \end{aligned} \quad (5.42)$$

Evaluating the revenue equation (2.4) at the monopolistically competitive price yields: $r(\varphi, \gamma_3) = (\rho P \gamma_3)^{\sigma-1} \varphi^{\sigma-1}$. Using this expression and equation (5.42), the distribution $F(r)$ for $r \geq r(\varphi_3, \gamma_3)$ is:

$$\begin{aligned} F(r) &= \frac{Pr((\rho P \gamma_3)^{\sigma-1} \varphi^{\sigma-1} \leq r) - Pr((\rho P \gamma_3)^{\sigma-1} \varphi^{\sigma-1} \leq (\rho P \gamma_3)^{\sigma-1} (\varphi_3)^{\sigma-1})}{Pr((\rho P \gamma_3)^{\sigma-1} \varphi^{\sigma-1} \geq (\rho P \gamma_3)^{\sigma-1} (\varphi_3)^{\sigma-1})}, \\ &= \frac{Pr\left(\varphi \leq \frac{r^{\frac{1}{\sigma-1}}}{\rho P \gamma_3}\right) - Pr(\varphi \leq \varphi_3)}{Pr(\varphi \geq \varphi_3)} = \int_{\varphi_3}^{\frac{r^{\frac{1}{\sigma-1}}}{\rho P \gamma_3}} \frac{g(\varphi)}{1 - G(\varphi_3)} d\varphi \end{aligned} \quad (5.43)$$

It follows that the CDF of value added for the subset of firms that have adopted steam power is:

$$F(r) = \begin{cases} 0 & \text{if } r < r(\varphi_3, \gamma_3) \\ 1 - (r(\varphi_3, \gamma_3))^{\frac{\theta}{\sigma-1}} r^{-\frac{\theta}{\sigma-1}} & \text{if } r \geq r(\varphi_3, \gamma_3) \end{cases} \quad (5.44)$$

Appendix B:

B.1. Proofs and Derivations from Chapter 3

Derivation of the policy functions $c_{i,t}^*$ and $q_{i,t+1}^*$

Starting from (3.20) and substituting the budget constraint into the period objective function yields:

$$V(q_{i,t}, \pi_{i,t}, g_t, P_t) = \max_{q_{i,t+1}} \ln((P_t + \pi_{i,t})q_{i,t}(1 + g_t) - P_t q_{i,t+1}) \\ + \frac{1-\lambda}{1+r} E_t V(q_{i,t+1}, \pi_{i,t+1}, g_{t+1}, P_{t+1}) + \frac{\lambda}{1+r} \left(\ln(\eta P_t q_{i,t+1}) + \frac{\ln(y)}{r} \right)$$

The first order condition with respect to $q_{i,t+1}$ is:

$$\frac{P_t}{(P_t + \pi_{i,t})q_{i,t}(1 + g_t) - P_t q_{i,t+1}} = \frac{1-\lambda}{1+r} E_t V'(q_{i,t+1}, \pi_{i,t+1}, g_{t+1}, P_{t+1}) + \frac{\lambda}{1+r} \frac{1}{q_{i,t+1}} \quad (5.45)$$

Under mild regularity conditions (see Stokey et al. (1989), chapter 4), one can apply the envelope theorem:

$$V'(q_{i,t}, \pi_{i,t}, g_t, P_t) = \frac{(P_t + \pi_{i,t})(1 + g_t)}{(P_t + \pi_{i,t})q_{i,t}(1 + g_t) - P_t q_{i,t+1}} \\ \Rightarrow V'(q_{i,t+1}, \pi_{i,t+1}, g_{t+1}, P_{t+1}) = \frac{(P_{t+1} + \pi_{i,t+1})(1 + g_{t+1})}{(P_{t+1} + \pi_{i,t+1})q_{i,t+1}(1 + g_{t+1}) - P_{t+1} q_{i,t+2}} \quad (5.46)$$

Substituting (5.46) into (5.45) and using the definitions of $c_{i,t}$ and $c_{i,t+1}$ yields the Euler equation:

$$\frac{P_t}{c_{i,t}} = \frac{1-\lambda}{1+r} E_t \left[\frac{P_{t+1} + \pi_{i,t+1}}{c_{i,t+1}} (1+g_{t+1}) \right] + \frac{\lambda}{1+r} \frac{1}{q_{i,t+1}}$$

Next I use the guess and verify method for solving for the optimal policy functions $c_{i,t}^*$ and $q_{i,t+1}^*$. My guess is as follows:

$$c_{i,t}^* = \alpha (P_t + \pi_{i,t}) q_{i,t} (1+g_t)$$

$$P_t q_{i,t+1}^* = (1-\alpha) (P_t + \pi_{i,t}) q_{i,t} (1+g_t)$$

Where my guess is that α is a constant. Using the Euler equation and my guess of $c_{i,t}^*$:

$$\begin{aligned} \frac{P_t}{\alpha (P_t + \pi_{i,t}) q_{i,t} (1+g_t)} &= \frac{1-\lambda}{1+r} E_t \left[\frac{(P_{t+1} + \pi_{i,t+1}) (1+g_{t+1})}{\alpha (P_{t+1} + \pi_{i,t+1}) q_{i,t+1} (1+g_{t+1})} \right] + \frac{\lambda}{1+r} \frac{1}{q_{i,t+1}} \\ \Rightarrow \frac{P_t q_{i,t+1}}{(P_t + \pi_{i,t}) q_{i,t} (1+g_t)} &= \frac{1-\lambda}{1+r} + \frac{\lambda}{1+r} \alpha \end{aligned}$$

Finally, implementing my guess of $P_t q_{i,t+1}^*$:

$$1-\alpha = \frac{1-\lambda}{1+r} + \frac{\lambda}{1+r} \alpha \quad \Rightarrow \quad 1-\alpha = \frac{1}{1+r+\lambda}; \quad \alpha = \frac{r+\lambda}{1+r+\lambda}$$

Thus verifying my guess that α is a constant. The optimal policy functions $c_{i,t}^*$ and $q_{i,t+1}^*$ are therefore:

$$c_{i,t}^* = \frac{r+\lambda}{1+r+\lambda} (P_t + \pi_{i,t}) q_{i,t} (1+g_t); \quad P_t q_{i,t+1}^* = \frac{1}{1+r+\lambda} (P_t + \pi_{i,t}) q_{i,t} (1+g_t)$$

Derivation of $V(q_{i,ss})$ and q_{ss}^* in the deterministic equilibrium

In order to analyze the effects of a price ceiling on the exit decisions of farmers, it is convenient to first derive the value function $V(q_{i,ss})$ and minimum viable scale of production q_{ss}^* as a function of P prior to substituting in the equilibrium value $P = \bar{\pi}/(r + \lambda)$. Under the conjecture that the value function is of the form $V(q_{i,ss}) = A + B \ln(q_{i,ss})$ and using the equilibrium values $q_{i,ss} = (1 + \bar{\pi}/P) q_{i,ss}/(1 + r + \lambda)$ and $c_{i,ss} = (P + \bar{\pi})q_{i,ss}(r + \lambda)/(1 + r + \lambda)$, equation (3.20) can be written:

$$A + B \ln(q_{i,ss}) = \ln \left(\frac{r + \lambda}{1 + r + \lambda} (P + \bar{\pi}) q_{i,ss} \right) + \frac{1 - \lambda}{1 + r} \left\{ A + B \ln \left(\frac{1}{1 + r + \lambda} \left(1 + \frac{\bar{\pi}}{P} \right) q_{i,ss} \right) \right\} \\ + \frac{\lambda}{1 + r} \left(\ln(\eta P q_{i,ss}) + \frac{\ln(y)}{r} \right)$$

The conjecture is verified by solving for A and B from the above equation, where $B = (1 + r + \lambda)/(r + \lambda)$ and

$$A = \frac{(1+r)(r+\lambda)+(1-\lambda)(1+r+\lambda)}{(r+\lambda)^2} \ln \left(\frac{P+\bar{\pi}}{1+r+\lambda} \right) - \frac{1+r-2\lambda(r+\lambda)}{(r+\lambda)^2} \ln(P) + \frac{\lambda}{r+\lambda} \left(\ln(\eta) + \frac{\ln(y)}{r} \right) + \frac{1+r}{r+\lambda} \ln(r + \lambda).$$

The representation of $V(q_{i,ss})$ in equation (3.25) can be derived by substituting $P = \bar{\pi}/(r + \lambda)$ into the above expression for A.

As discussed in section 3, q_{ss}^* is defined by the equality $V(q_{i,ss}) = V^e(q_{i,ss})$. Solving this equality for q_{ss}^* yields:

$$q_{ss}^* = y \eta^{-\lambda} (P + \bar{\pi})^{-\frac{(1-\lambda)(1+2(r+\lambda))}{r+\lambda}} P^{\frac{1+r-2\lambda(r+\lambda)}{r+\lambda}} (r + \lambda)^{-(1+r)} (1 + r + \lambda)^{\frac{(1+r)(r+\lambda)+(1-\lambda)(1+r+\lambda)}{r+\lambda}} \quad (5.47)$$

The representation of q_{ss}^* in equation (3.26) can be derived by substituting $P = \bar{\pi}/(r + \lambda)$ into equation (5.47). Equation (5.47) is a convenient representation of q_{ss}^* as it can be used to analyze the effect a marginal change in the P , holding constant

$\bar{\pi}$, as is the case with price ceiling legislation. Evaluating the partial derivative of q_{ss}^* with respect to P at the equilibrium price $P = \bar{\pi}/(r + \lambda)$ yields:

$$\frac{\partial q_{ss}^*}{\partial P} \Big|_{P=\frac{\bar{\pi}}{r+\lambda}} = \frac{y(r + \lambda)^{1-r}}{\bar{\pi}^2 \eta \lambda (1 + r + \lambda)^{1-r-\lambda}} (r - 2\lambda(r + \lambda)) \quad (5.48)$$

This derivative is indeterminate of sign as the final term may be positive or negative depending on the values of r and λ . In section 4.2 I estimated that farmers' perception of policy risk was 4.63% between 1999-2005. In that section I assumed that farmers' discount rate during this period was equal to the prime interest rate plus two percent, 7.262%. The partial derivative in (5.48) is positive when evaluated at these values.