THREE ESSAYS IN ECONOMICS

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A Thesis Submitted to the School of Graduate Studies in Partial Fulfilment of the Requirements for the Degree Doctor of Philosophy

McMaster University DOCTOR OF PHILOSOPHY (2019) Hamilton, Ontario (Economics)
TITLE: Three Essays in Economics
AUTHOR: Md Mahbubur Rahman
SUPERVISOR: Professor Alok Johri
NUMBER OF PAGES: xi, 108

Abstract

This thesis attempts to answer three important questions: 1) Why did India's relative price of investment rise in 80s and fall in 1990s and afterwards? 2) Why is agricultural productivity very low in India? and 3) Did the pro-natalist policy in Quebec accomplish its goal of increasing fertility? Specifically, this thesis comprises of three essays. Chapter 1 builds a simple dynamic general equilibrium model calibrated to Indian data, in order to explore the impact of capital import substitution policies and their reform post-1991. The model delivers a 23% rise before reform and a 28% fall thereafter. Chapter 2 develops a tractable quantitative framework by incorporating one potential explanation - if residing in a village provides access to a network that effectively insures against income fluctuations, then households are less willing to live in cities where labor income risk is uninsured. This chapter shows that implementation of a social insurance system in the urban area could have raised the labor productivity agricultural sector. Chapter 3 studies the effects of a pro-natalist policy in Quebec and finds Quebec's baby bonus accomplished its goal of increasing fertility. It finds a large response for third and higher-order births for which the bonus was more generous. Interestingly, it also finds a stronger response if there were two previous sons or a previous son and daughter rather than two previous daughters.

Acknowledgements

I am immensely indebted to my PhD supervisors Professor Alok Johri for his invaluable support and guidance that have made it possible for me to complete my dissertation. I am also indebted to Professor Marc-André Letendre, Professor Pau S. Pujolàs for their continuous time and energy without fail. I would like to express my greatest appreciation to Oksana M. Leukhina for her support and guidance.

My special thanks go to Byron G. Spencer, Bettina Brueggemann, Gajendran Raveendranathan and Zachary Mahone for elevating me and giving me the strongest platform possible for my growth. I would also like to thank Arthur Sweetman, Kate Cuff, Philip DeCicca, Jeffrey S. Racine, Michel Grignon, Svetlana Demidova and Michael Veall, all of whom have played a significant role in my PhD journey. Also, I would like to thank my co-authors Natalie Malak and Terry A. Yip for their patience and am looking forward to many collaborations in the future.

I thank the friends I made during my graduate study, Anthony Hong, Bradley Minaker, Camille Simardone, Farhana Khanam, Felix Li, Grant Gibson, Helen Zhang, Ionut (Johnny) Cotoc, James Spencer, Joel Wanger, Josip Lesica, Julius Owusu, Karen Ugarte Bravo, Lan Wei, Luc Clair, Muhebullah Karimzada, N. T. Khuong Truong, Nadine Chami, Samir Mainuddin, Sabreena Obaid, Saeed Rana, Sean Sexton, Shahed Kaiser Khan, Shaun Shaikh, Stephanie Houle, Stephanie Thomas, Wei Yang and Zvezdomir Todorov. Additionally, the staff in the Economics department have been nothing short of exceptional. I would like to thank the analysts at the McMaster RDC, Peter Kitchen and Mustafa Ornek for their help.

Finally, I would like to thank my family and friends. They are always beside me in low and high and continue to be the inspiration in every pursuit in my life. My parents' and grandparent' love and encouragement have allowed me to pursue my academic endeavors. My wife, Munima Siddika, who would follow me to the ends of the earth, has been my biggest supporter. I would also like to thank my little daughter Arissa for her love.

This thesis is dedicated to my mother Rokeya Sultana.

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Declaration of Academic Achievement

Chapter 1 is co-authored with Professor Alok Johri. Chapter 2 is co-authored with Oksana M. Leukhina. Chapter 3 is co-authored with Natalie Malak and Terry A. Yip. I participated in all stages of the research.

Introduction

Some key questions in economics form the underlying motivation for this thesis. First, what determines aggregate labor productivity in developing countries and how it could be improved through policy? Second, do pro-natalist policies work in a developed economy, who is taking advantage of the incentives provided, and how costly are the programs? Specifically, this thesis comprises of three essays.

In the first chapter of my thesis, "The Rise and Fall of India's Relative Investment Price: A Tale of Policy Error and Reform", we study the puzzling dynamics of the relative price of capital goods in India. The relative price of capital rose by 44 percent in the 80s in India and then it declined by 26 percent since the economic reforms implemented in 1991. What we saw in the 80s is difficult to match with empirical evidence from other countries. Empirically what we see in other countries is that the relative price of capital declines with economic growth. India has experienced GDP growth over the same time period. So, the price behavior in India is very atypical in the 1980s and I investigate why this is the case. It turns out that throughout the 80s India maintained a very strict import policy. All capital importers needed licenses and were subjected to a strict import quantity restriction. In 1991, the newly elected government removed these restrictions. This change in policy coincided with the trend shift in the capital price. We build a general equilibrium model that allows me to quantitatively analyze the effect of trade policy on the rise and fall of the price of capital. We find that the presence of the restrictive trade policy in the 80s in the face of productivity growth accounted for 50 percent of the observed rise in the relative price of capital, while the removal of quantity restrictions accounted fully for the observed decline.

In Chapter 2, "Why is Agricultural Productivity So Low in Poor Countries? – The Case of India", we try to understand why labor productivity in the agricultural sector is so low in poor countries. In India, for example, labor productivity in agriculture is only one fifth of the urban labor productivity. We propose and investigate the following potential explanation of this productivity gap. In villages, people have access to a network that effectively insures them against income fluctuations whereas in cities there are limited opportunities to insure against labor income risk. This means that households are less willing to live in cities and find villages more attractive. As a result, labor remains abundant and therefore cheap in rural areas. Farmers incentives to mechanize production remain weak. In other words, it is optimal for them to choose primitive ways of production that rely more on labor and less on capital. The size of the farm remains small and labor productivity in the agricultural sector remains low. In order to understand the quantitative importance of this mechanism, we calibrate the model to Indian data and study an abstract policy intervention: a provision of complete insurance against earnings risk in the city. The policy intervention increases labor productivity in agricultural sector by 37 percent and reduces the urban-rural productivity gap by 30 percent. This effect comes about because of the 7 percent drop in agricultural share of employment, which encourages an inflow of capital in the agricultural sector and raises the average farm size by 12 percent.

In Chapter 3, "Baby Bonus, Anyone? Examining Heterogeneous Responses to a Pro-Natalist Policy", we examine the impact of the Allowance for Newborn Children, a universal baby bonus offered by the Canadian province of Quebec, on birth order, sibship sex composition, income, and education.[1] We find a large response for third and higher-order births for which the bonus was more generous. Interestingly, though, we find stronger response if there were two previous sons or a previous son and daughter rather than two previous daughters. We also find, in addition to a transitory effect, a permanent effect, with the greatest increase in one daughter-two son families among three-child households. Moreover, we find a hump shape response by income group, with the greatest response from middle-income families. Also, women with at least some post-secondary education respond more to the policy than those with less. These findings suggest that properly structured pro-natal policies can successfully increase fertility among different segments of the population while simultaneously diminishing the effect of gender preferences and fertility disparity related to women's education.

[1] This paper will be published in the Journal of Population Economics in October 2019. The citation of the paper is: Malak, Natalie, Md Mahbubur Rahman, and Terry A. Yip. "Baby Bonus, Anyone? Examining Heterogeneous Responses to a Pro-natalist Policy." *Journal of Population Economics*, 2019, 32(4), 1205–1246.

Chapter 1

The Rise and Fall of India's Relative Investment Price: A Tale of Policy Error and Reform

1.1 Introduction

The behavior of the relative price of investment in India stands in fascinating contrast to its well known fall in the US in recent decades (Greenwood et al., 1997).[1] As seen in Figure 1.1, relative to the Penn World Table benchmark index (Feenstra et al., 2015), the relative price of investment in India rose 44 percent from 1981 to 1991 and subsequently fell 26 percent from 1991 to 2006. To contextualize the magnitude of this change in relative price, we can look at cross country differences in the relative price of investment in 1991. The average value of the relative price of investment for G7 nations in 1991 was 0.88 while the average for all other nations was 1.38 which is 57 percent higher. Similarly the one decade rise seen in India of 44 percent is equivalent to moving from the United States to nations such as Antigua and Barbuda or Estonia in terms of percent difference in the relative price of investment.

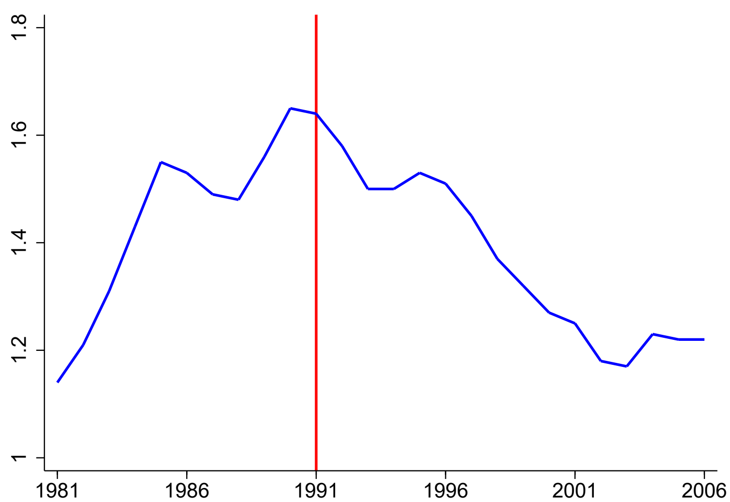

Figure 1.1: The relative price of investment in India from 1981 to 2006

Note: The vertical line denotes the year 1991 when capital import reform begins. *Source:* The Penn World Table 9.0.

The sudden change in direction in the relative price of investment in India is tantalizingly coincident with a period of rapid economic reform in India and the concomitant increase in the growth rate of Indian GDP. These observations raise a

[1]The decrease in the relative price of investment can also been seen in world indexes starting in the early 1980s (Karabarbounis and Neiman, 2014).

number of questions that we seek to answer in this paper. Why did the relative price of investment rise in India during the 1980s while it fell in the developed world? Did the sudden change in direction in 1991 have something to do with the unexpected change in policies instituted by the Indian state during the reform period beginning in 1991 and beyond?[2] If so, what was the contribution of these policy shifts to the increase in the growth rate of GDP experienced by India over the next decade and a half?

When thinking about the divergent paths of the relative price of investment goods in India and the world benchmark, it is natural to focus on policy distortions specific to the import of machines into India. This is especially true when the vast majority of capital goods are produced in a few developed nations (Mutreja et al., 2016). These distortions were large and came from several sources. Before reforms, capital good imports into India faced very high tariff rates — the United Nations Conference on Trade and Development (UNCTAD) calculates the weighted average tariff rate on capital to be 72.7 percent in 1990. Similarly Hasan et al. (2007) report that in 1988 electrical machinery faced a tariff rate of 143 percent, transport equipment 130 percent, and other machinery 140 percent approximately. In addition, there existed pervasive non-tariff barriers on the import of capital goods which required import licenses to be obtained from the government. Hasan et al. (2007) report that quantitative restrictions applied to 90 percent of the value added in manufacturing. The coverage rates of non-tariff barriers for the import of goods in the machinery category was 77 percent, while it was 79 percent for the electrical machinery category and an even higher 82 percent for the import of transport equipment. After 1990, tariff rates on imported capital goods fell from a weighted average of 72.7 percent to 7.6 percent by 2006 in a series of steps. Moreover import licensing was removed from a number of capital good categories that quickly expanded such that they became freely importable by 1993.[3]

In this paper we argue that both the rise in the relative price of investment in the pre-reform period and the fall during the reform period were closely linked to Indian capital import substitution policies and their removal after 1991. The fall in prices seen over the reform period are relatively easy to understand. We would expect the price of investment goods to fall as tariffs on imported machines used in the production

[2] Most observers agree that the change in policy regime in India around 1991 was unexpected. For example, see Goyal (1996) and Goldberg et al. (2010).

[3] Additional figures and details can be found in Kotwal et al. (2011) and the references therein.

of these goods are reduced over time.[4] The rise in the pre-reform decade is a little less obvious. Our explanation is based on the insight that any decade-long endogenous rise in the relative price of investment must come from an increasing relative scarcity of machines. The import substitution policies instituted by the Indian state provide a well documented source of scarcity. Import license requirements were put in place precisely to restrict the amount of capital goods that could be brought into India. This policy-induced scarcity created a wedge between domestic and world prices of capital good imports, over and above the already high tariff rates. Indeed policy makers must have found that the tariff barriers were insufficient to squelch the demand for foreign machines and thus resorted to outright quantity restrictions in the form of licenses. This relative scarcity increased with time as the demand for imported capital goods grew along with the economy in the 1980s, both due to productivity growth (which picked up over this period) and due to rapid population growth while the supply was kept relatively tight by policy. The impact of these policies can be seen in a decline in the ratio of capital imports to output which fell by 28 percent between 1981 and 1991.[5]

In order to quantitatively explore the contribution of these import substitution policies to the rise and fall of the relative price of investment in India, we build a simple dynamic general equilibrium model of a small open economy in which foreign capital goods are an input into the production of domestic investment goods. In the pre-reform period, the import of capital goods is capped at a fixed amount each period as specified by the government so that the domestic price paid by firms for foreign capital goods is determined by a market clearing domestic price which can differ from the world price. As demand for foreign capital goods rises, the domestic price of foreign capital goods rises as well, even in the presence of constant tariff rates because the cap on imported capital goods falls further and further behind demand. During the reform period, this cap is removed so that capital goods can be freely imported into the economy, once a tariff is paid to the government. In the absence of a constraint on imports, the wedge between world and domestic prices of foreign capital goods

[4]A number of studies have established the positive impact of Indian tariff reductions on firm choices and performance, though we are not aware of any that study the impact on the relative price of investment. For example, Topalova and Khandelwal (2011) estimate that the sudden fall in import tariff rates in 1991 had a positive impact on firm level productivity of domestic manufacturing firms. Goldberg et al. (2010) and Goldberg et al. (2009) find that the decline in trade tariff rates led to a large expansion in new products and imported input use by Indian firms. Bollard et al. (2013) however find that reforms have only a limited ability to explain TFP growth in large existing firms. See also Chamarbagwala and Sharma (2011).

[5]Details about the construction of the capital import share can be found in section 1.3.1 below.

becomes exogenous and is solely determined by tariff rates. We calibrate this model to Indian macroeconomic data and explore two scenarios. First, in the pre-reform phase, we ask, how much does the relative price of investment increase in the model when we embed the actual growth in productivity and employment experienced by India from 1981 to 1991.[6] Next we use observed reductions to tariff rates on imported capital goods from the UNCTAD's Trade Analysis & Information System (TRAINS) database and calculate the implied fall in the relative price of investment in the calibrated model without an import constraint. Our results suggest that we can generate a rise of about 23 percent in the relative price of investment before reforms begin, and thereafter, a decline of roughly 28 percent. This large decline comes from both the removal of quantity restrictions as well as the decline in tariff rates. These movements in the relative price of investment have great significance for the economy. The model implies that GDP per worker was 2.9 percent lower in 1991 compared to a decade earlier purely due to the rising distortion caused by import restrictions on capital goods. In addition, the 64 percentage point reduction in capital import tariff rates and removal of quantity restrictions raised GDP per worker permanently by 17.8 percent. Turning to the transitional dynamics induced by the reduction in capital import tariff rates, they alone account for one fifth of the rise in the growth rate of GDP per worker observed between 1991 and 2006. Bosworth and Collins (2008) report that output per worker growth almost doubled from 2.4 to 4.6 percent per annum in the reform period. Consistent with our story, the authors report that the contribution of physical capital also doubled from 0.9 percent to 1.8 percent.

An interesting feature of our model is an endogenously rising policy distortion in the pre-reform period which increases with the overall size of the economy.[7] To our knowledge, this is the first paper that provides an endogenous explanation for medium term movement in the relative price of investment over time. Our work is related to the literature that explains cross-country differences in the relative price of investment based on exogenous relative productivity differences in the investment versus consumption sector (Hsieh and Klenow, 2007) and especially to exogenous differences in investment distortions. For example Restuccia and Urrutia (2001) establish the large

[6] We start our analysis in the 1980's because the previous decade was tumultuous, marked by the 1971 war with Pakistan, the OPEC oil price shock of 1973 and the political crisis known as the emergency from 1975 to 1977. Perhaps due to these disturbances, output per worker barely grew in the 1970s and TFP growth was negative (see Rodrik and Subramanian (2005)). Kochhar et al. (2006) also begin their analysis of Indian development in the 1980s.

[7] The model also contains exogenous sources of distortion which are discussed more fully later.

dispersion in the relative price of investment across countries and use an exogenous stochastic process for distortions to investment to account for these facts.[8] Our work differs from these in two ways. First we do not focus on cross-sectional differences in the level of relative prices in, say, India and the United States at a point in time. Instead we explain why the relative price of investment in India increased compared to US for a long period of time. Second, the change in the relative price of investment in our model is driven by a distortion that grows over time with the economy because of a policy induced scarcity of foreign machines. As a result the model has no sectoral differences in productivity trends. To the extent that there were differences in trends in productivity in investment versus consumption good sectors, these would complement our quantitative results. Indeed an additional 21 percent rise in India's relative price of investment remains unaccounted by our calibrated model. Our interest in medium term trends in the relative price of investment is shared by Karabarbounis and Neiman (2014) which links the fall in the relative price of investment to the recent global fall in labor share; however, changes in the relative price of investment are driven by exogenous shocks to the productivity of the investment sector in that study. We share an interest in exploring the quantitative implications of population growth on development with Leukhina and Turnovsky (2016) who study the transition from agriculture to manufacturing in England.

The rest of the paper is structured as follows. Section 1.2 presents the model while section 1.3 discusses the data used in the study as well as the calibration of model parameters. Section 1.4 presents quantitative results from the benchmark calibration as well as some sensitivity analysis. Concluding remarks are followed by an appendix that outlines the solution methods used in our paper.

1.2 Model

We model a standard small open economy that imports a capital good at a given world price and combines it with a domestic final good to create the domestic investment good used for capital accumulation. There are three type of firms in the economy: a representative final good producer, a representative investment good producer and

[8]Our model is also related to the literature that links the relative price of investment to growth, investment, income and productivity differences across nations. Jones (1994) provides an early link between the relative price of investment and economic growth. See also Armenter and Lahiri (2012) and Restuccia (2004).

a representative importer that faces a capital import restriction. All firms behave competitively. In addition, the model has a representative household and government.

1.2.1 The Household's Problem

The benevolent head of an infinitely lived representative household of size L_t obtains utility from sequences of total consumption, C_t, of the final good with lifetime utility defined as

$$\mathcal{U} = \sum_{t=0}^{\infty} \beta^t \log C_t \tag{1.1}$$

where β, $0 < \beta < 1$, is the household's subjective discount factor.[9]

The household supplies one unit of labor per person so that it supplies L_t units of total labor inelastically to the final good producer. In each period it earns a wage equal to w_t per unit of labor. In addition it earns capital income by renting out its capital stock K_t at the rental rate r_t and also receives profits from the final good producer, the investment good producer, the importer and receives lump-sum transfers, T_t, from the government. At the end of each period, the household chooses its total consumption, C_t, (divided equally among members), and buys domestic investment good, I_t, at price, q_t, which will be our notation for the relative price of investment. All prices are expressed in units of the final good. The household budget constraint is

$$C_t + q_t I_t = w_t L_t + r_t K_t + \Pi_t^{\mathrm{Y}} + \Pi_t^{\mathrm{I}} + \Pi_t^{\mathrm{imp}} + T_t \tag{1.2}$$

and the law of motion for capital is

$$K_{t+1} = (1-\delta)K_t + I_t \tag{1.3}$$

where δ is the depreciation rate of capital. The household chooses sequences of C_t, K_{t+1} to maximize (1.1) subject to (1.2) and (1.3), and the initial condition, $K_0 > 0$, which yields the first order condition:

$$\frac{q_t}{C_t} = \beta \frac{1}{C_{t+1}} \left(r_{t+1} + q_{t+1}(1-\delta) \right) \tag{1.4}$$

[9]We follow the literature in the use of log preferences. See Restuccia (2004) for example.

1.2.2 The Final Good Firm's Problem

The perfectly competitive final good producer operates a constant returns to scale technology given by

$$Y_t = K_t^{\alpha}(Z_t L_t)^{1-\alpha} \tag{1.5}$$

where Z_t is productivity that grows at the exogenous rate γ_z. Our notation presupposes market clearing in factor markets, therefore we do not distinguish between quantities supplied and demanded in these markets. As such, since each member of the household inelastically supplies one unit of labor, L_t, measures, not only the hours hired by the firm, but also the size of the working population which grows exogenously according to (1.6)

$$L_t = \gamma_l L_{t-1}. \tag{1.6}$$

The firm sells its output in the final good market to the household for consumption, and to the investment good producer as an input in investment good production. Standard efficiency conditions for the producer are omitted for brevity.

1.2.3 The Investment Good Firm's Problem

The representative investment good producer combines units of the imported capital good with units of the domestic final good to produce the domestic investment goods using the following technology :

$$I_t = D_t^{\eta} M_t^{1-\eta} \tag{1.7}$$

where D_t refers to units of the domestic final good, and M_t to units of the imported capital good purchased. Domestic and foreign capital goods are usually combined using the Cobb-Douglas specification in the literature (see Boileau (2002) and Hsieh and Klenow (2007) for example).[10] The firm buys M_t from the importer at price p_t^m, and sells the produced investment good, I_t to household at price q_t. The investment good producer chooses D_t and M_t to maximize its profits given by

$$\Pi_t^{\mathrm{I}} = q_t I_t - D_t - p_t^m M_t, \tag{1.8}$$

[10] Robustness analysis is conducted with constant elasticity of substitution technology, see section 1.4.3.

yielding the first order conditions:

$$\eta q_t D_t^{\eta-1} M_t^{1-\eta} = 1 \tag{1.9}$$

$$(1-\eta) q_t D_t^{\eta} M_t^{-\eta} = p_t^m \tag{1.10}$$

Combining equations (1.9) and (1.10), we get a relationship between the intensity of imported capital use in the economy and the domestic price of imported capital goods:

$$\frac{D_t}{M_t} = \frac{\eta}{1-\eta} p_t^m \tag{1.11}$$

which shows that the government can pursue its pre-reform agenda of import substitution by implementing policies that inflate p_t^m. We also get the following relationship between q_t and p_t^m,

$$q_t = \frac{1}{\eta} \left(\frac{\eta p_t^m}{1-\eta} \right)^{1-\eta} \tag{1.12}$$

which further clarifies the mechanism by which our model will operate to influence the relative price of investment over time.

1.2.4 The Government

The government plays a limited role in our model. It follows capital import substitution policies by imposing a tariff, θ_t, on each unit of imported capital goods. This is paid by the importer. We assume the government runs a balanced budget so that all revenues from the import tariff are rebated to the household as a lump-sum transfer, T_t. In addition we interpret the license requirements on imported capital goods as a government set capital import limit, $\bar{M}_t$, that potentially changes with time. We assume the importing firm must obtain one license per unit of imported capital goods so that we can think of the domestic market for imported capital in symmetry with the market for licenses. The importing firm is restricted to importing no more than $\bar{M}_t$ units of capital goods into the country in the pre-reform period whereas there is no constraint on the importer after reforms begin. In practice import licenses were often expressed in nominal terms. This meant that periods of rapid depreciation of the Rupee inadvertently made the capital import limit even tighter in physical units. We discuss this issue in more detail when calibrating the path of $\bar{M}_t$.

1.2.5 The Importing Firm's Problem

The representative importing firm brings foreign capital goods, M_t, from outside the country, taking as given the world price p^w. In addition it must pay the tariff, θ_t, to the government. The importer then sells M_t units of imported capital goods at the market clearing price, p_t^m. The importer's profits are given by

$$\Pi_t^{\text{imp}} = p_t^m M_t - p^w(1+\theta_t)M_t \tag{1.13}$$

In the reform period, when the import limit is effectively infinite, the importer maximizes profits by choosing M_t. Before reforms, if the constraint imposed by the government binds, $M_t = \bar{M}_t$, otherwise it is chosen to maximize profits. Efficiency conditions imply that $p_t^m \geq p^w(1+\theta_t)$. When the domestic price of imported capital goods exceeds the cost to the importer, we assume profits are repatriated to the household in a lump sum fashion. After reforms begin, the importer makes zero profits. In India, imports of machines and other inputs were often carried out by central government agencies such that any profits earned flowed into the coffers of the government. Since both tariff revenue and profits from imports flow back to the household, we could easily have pooled the importer into the government without any loss of results. Note that the premium charged by the importer over and above the tariff inclusive price could also be interpreted in terms of bribes paid to bureaucrats in order to obtain a license to import, where the bribe amount is determined by supply of and demand for licenses. Note also that the price, p_t^m, which measures the degree of distortion in the domestic market for foreign capital goods, is determined endogenously in equilibrium in the pre-reform period. During the reform period p_t^m is the sum of two exogenous components, namely the tariff rate and the world price of imported capital goods. The world price of imported capital goods is held constant in the model because we want to generate movements in the relative price of investment in India relative to the world benchmark index for the relative price of investment in the Penn World Table. This ensures that any movement in the relative price solely emerges from domestic sources in India in the model.

1.2.6 Equilibrium

Definition: Given the initial conditions, the equilibrium of this economy is given by sequences of C_t, I_t, K_{t+1}, D_t, M_t, T_t and prices w_t, r_t, q_t, p^w and p_t^m where $t = \{0, \ldots, \infty\}$ such that (i) given w_t, r_t and q_t the representative household chooses C_t, I_t, K_{t+1} to solve its utility maximization problem using (1.4); (ii) given w_t and r_t, the final good producing firm chooses K_t, L_t to solve its profit maximization problem; (iii) given p_t^m, the investment good producing firm chooses D_t and M_t to solve its profit maximization problem using (1.9) and (1.10); (iv) given p^w, θ_t and the government imposed restriction $\bar{M}_t$, the importer chooses M_t to solve its problem using (1.13) and; (v) markets for labor, capital, investment goods, foreign capital goods and final goods clear; (vi) the government budget is balanced; and (vii) the aggregate resource constraint, $C_t + D_t = Y_t - p^w M_t$, holds.

At this point it may be useful to discuss the dynamics of the model in two situations, when $\bar{M}_t$ is binding and when it is not. We begin with the latter situation. When $\bar{M}_t$ is not binding, p_t^m is effectively exogenous and only responds to changes in tariff rates. From (1.12) we can see that in this situation the relative price of investment, q_t, is constant and the economy follows a balanced growth path where all other variables grow at a constant rate given by growth of productivity and employment. When the import constraint is binding, there are two possible scenarios. In the first scenario, $\bar{M}_t$ grows at the same rate as productivity and employment growth, therefore, q_t is still constant and the economy follows a balanced growth path. In the second scenario, the import constraint grows at a slower rate, then p_t^m and q_t both rise and the economy is no longer on a balanced growth path. This occurs because the rise in the price of imported capital goods causes the investment producer to change the optimal mix of domestic and foreign goods used in the production of investment goods. As a result $\frac{D_t}{M_t}$ falls over time as is clear from (1.11). Our solution method, discussed in section 1.4.1, provides a terminal period for this scenario after which the economy returns to a balanced growth path.

1.3 Data Definitions and Calibration

In this section we describe the data used in our study and discuss how the parameters of our model were chosen.

1.3.1 Data

The Penn World Table 9.0 (PWT) (Feenstra et al., 2015) provides data on relative levels of income, output, inputs and productivity in 182 countries between 1950 and 2014. Below, the series from the PWT used in our paper are discussed with the series name in parentheses. The price of consumption (‘pl_c’) and the price of investment (‘pl_i’) are constructed using both a purchasing power measure and a “reference price” (which we refer to as the world benchmark). The reference price is calculated using the *quantity-weighted average over countries of prices of each good*. The relative price of investment is constructed by taking the ratio of the price of investment and the price of consumption. The aggregate depreciation rate (‘delta’) is a weighted average of the following categories: structures (residential and non-residential), transport equipment, computers, communication equipment, software and other machinery and assets. To calculate the capital import share, we use import share data measured in current purchasing power parity units on the following categories of merchandise trade: industrial supplies (‘csh_m2’), fuels and lubricants (‘csh_m3’), capital goods (‘csh_m4’), and transport equipment (‘csh_m5’).[11] These import shares are measured as the ratio of import expenditure by category to nominal GDP at current prices and therefore contain movements in the prices of imports relative to the GDP deflater. To remove these prices, we construct the ratio of capital imports to consumption using:

$$\sum_{i=2}^{5} \frac{csh\_mi}{csh\_c} \times \frac{pl\_c}{pl\_mi}$$

- where (‘csh_c’) is the consumption share, the various import shares are: industrial supplies (‘csh_m2’), fuels and lubricants (‘csh_m3’), capital goods (‘csh_m4’), and transport equipment (‘csh_m5’) with corresponding import prices (‘pl_mi’), i=2,3,4,5.

To calculate productivity, Z_t, we use real GDP at constant 2011 national prices (‘rgdpna’), real capital stock at constant 2011 national prices (‘rkna’), number of persons engaged (‘emp’). To calculate the consumption-output ratio, we use the series, ‘Real consumption at constant 2011 national prices’ and divide it by ‘Real GDP at constant 2011 national prices’.

The UNCTAD’s Trade Analysis & Information System (TRAINS) database provides

[11] Imported capital categories used in our calibration are also similar to EU-KLEMS database (Jäger, 2016).

data on the average tariff rate (UNCTAD method) on capital goods (UNCTAD-SoP4 – Capital goods[12]).[13] This series is available from 1990.

1.3.2 Parameters

Our model is calibrated to match several features of the Indian economy in 1981 which is the starting year of our analysis. These parameters values are provided in Table 1.1. We assume that the economy was on a balanced growth path in 1981. Some evidence in support of this assumption can be seen in the relatively stable investment to output ratio during the decade of the 1970s when this ratio had an average value of 0.20. By contrast the investment output ratio declined in the 1980s and then rose in the 1990s displaying average values of 0.19 and 0.22 respectively (the investment output ratio is obtained from the PWT using the series 'csh_i').

Table 1.1: Parameters : benchmark calibration.

Parameter		Value	Source
Capital share in final good production	α	0.33	standard
Discount factor	β	0.9255	calibrated
Employment growth	γ_l	1.0326	PWT
Depreciation rate	δ	0.05	PWT
Productivity growth	γ_z	1.0212	PWT
Import share in investment good production	$(1-\eta)$	0.2650	calibrated
Tariff	θ	0.72 & 0.076	UNCTAD's TRAINS

There are two sources of growth in the model, the number of employed people, L_t, and the level of labor augmenting productivity, Z_t. We set the gross growth rate of the labor input, $\gamma_l = 1.0326$ to match the annualized growth rate of the employed population in India between 1981 and 2006. We also set the gross growth rate of productivity, $\gamma_z = 1.0212$ to match the observed growth in the labor augmenting productivity for India during the same period. Since the initial level of L_t and Z_t has no impact on the percentage change in q_t, we normalize the initial values of both to unity. The value of the depreciation rate, δ, in the capital accumulation equation is 0.05, which is obtained from the average of the annual reported value in the PWT.

[12]UNCTAD-SoP4 is a Harmonized System (HS) classification for capital goods.

[13]Data is accessed through the World Bank: WITS application; see http://wits.worldbank.org.

We calibrate β, the time preference parameter in the household utility function, and η, the share of domestic final goods in the production of investment goods, jointly using the consumption-output ratio in 1981 which is 0.84 and the capital import to consumption ratio in 1981 which is 0.03 in the PWT. We described the construction of this measure in the previous subsection. Turning to the final good production technology, we follow the literature in assuming a constant returns to scale production function of the Cobb-Douglas form. The capital share parameter, α, is set to 0.33, a standard value in the literature (Hsieh and Klenow (2007)) which is also close to the average value seen in this period.

Table 1.2: Tariff rates on capital imports.

Year	Tariff rate (%)	Step
1990	72.72	-
1992	52.62	Step 1
1996	29.30	Step 2
1997	21.24	Step 2
1999	26.60	Step 2
2000	21.70	Step 2
2001	22.37	Step 2
2002	20.79	Step 2
2003	19.65	Step 2
2004	21.73	Step 2
2005	9.60	Step 3
2006	7.06	Step 3

Note: Tariff rates are calculated using UNCTAD's averaging method. *Source:* UNCTAD's TRAINS database.

The pre-reform period differs crucially from the reform period due to the presence of the import constraint captured by $\bar{M}_t$. Obviously the path of this variable has a strong influence on the domestic price of imported capital goods, p_t^m, and through that on the level of the relative price of investment, q_t. Our approach for disciplining the quantitative analysis is to choose the most conservative level for $\bar{M}_t$ in 1981 while ensuring the constraint is actually binding in equilibrium. This level of imports is obtained when the 1981 price of imported capital goods, p_t^m, is equal to 1.72 which is composed of the world price of capital goods, $p^w = 1$, and the tariff rate of 72 percent. This is equivalent to a situation where the effective distortion caused by the non-tariff barrier is no greater than the actual tariff rate observed in 1990 (see Table 1.2). Note that our analysis focuses on the percentage change in the relative price of investment and not on the level in any year. The rise in the relative price discussed in the next section is not sensitive to the initial level of $\bar{M}_t$. We demonstrate this by recalculating the change in q_t for a lower value of $\bar{M}_t$ which implies an imported capital good price, $p_t^m = 2$ as opposed to the benchmark value of 1.72.

In order to discipline the path of $\bar{M}_t$ beyond the initial period we use the ratio of capital imports to consumption obtained from the PWT in 1991 which has a value of 0.014. In our model economy this ratio shrinks for three reasons: labor productivity growth, employment growth and a change in $\bar{M}_t$. As discussed above, we measure the growth rate of the former two factors directly from the PWT data. The third factor, which governs the increase or decrease of the capital import limit is not directly observed but can be extracted from model simulations. We pick the growth rate of $\bar{M}_t$ between 1981 and 1991 to be such that the ratio of capital imports to consumption is exactly equal to the data value in 1991 and 1981. Our simulations reveal that $\bar{M}_t$ must shrink at a rate of 3.5 percent each year to meet our target. In order to understand this tightening of import limits one needs to remember that the Rupee depreciated dramatically in the 1980's and this caused nominal import limits expressed in Rupees to shrink in real terms. Our finding of an aggregate tightening of 3.5 percent per annum should be viewed as the net effect of the real depreciation of the Rupee and a slowly liberalizing import policy which was not keeping pace with the falling value of the currency and the rising demand for imported capital goods due to population and productivity growth.

The pre-reform tariff rate, θ_t, on imported capital goods is set to 0.72. We use weighted average tariff rates on capital imports from 1991 until 2006, when θ_t falls to 0.076. Since these tariff rates are not available for every year, we try to capture the falling trend in rates by pooling the reductions into three discrete steps. We provide a table with the actual measure of the weighted tariff rate by year and our steps in Table 1.2. We stop our analysis of the reform period in 2006 to avoid the impact of of the financial crisis and the US trade collapse (see Ahn et al. (2011) for a discussion of the size of the collapse in world trade). Since our measure of the relative price of investment in India is calculated relative to a benchmark relative price of investment for the world, we normalize the world price of imported capital goods to unity.

1.4 Results

1.4.1 Before Reform

In this section, our goal is to get a quantitative sense of the ability of our calibrated model to produce a rise in the relative price of investment while also obtaining measures of the impact of the import substitution policy on output per worker. We begin our analysis by assuming that the economy is on a balanced growth path until 1981. In order to implement our solution method, we divide all growing variables by effective labor. For example, we define output per effective unit of labor as $y_t = \frac{Y_t}{Z_t L_t}$ and similarly for other variables, giving us an initial steady state in the transformed system. Since $\frac{\bar{M}_t}{Z_t L_t}$ shrinks every period after 1981, the economy is no longer in steady state. To solve the model we assume that agents expect the current policy on tariff rates and a tightening import constraint to remain in place for 50 years after which the government adjusts the import limit to keep pace with productivity and employment growth forever.[14] We compute the transition of the economy from 1981 to 1991 at which point an unexpected change in policy occurs.

We confirm that the pre-reform transition path of the economy is not significantly affected by our choice of terminal year for when the import limit stops shrinking. We obtained similar results with a 20, 50 and 100 years transition since there is little change in the path in the first 10 years. To solve the transition in the pre-reform period, we use the relaxation algorithm for a system of non-linear equations using the forward-looking method proposed by Boucekkine (1995). Key aspects of this algorithm involve a known initial and terminal condition and perfect foresight for agents regarding the path of exogenous variables.

Compared to 1981, the simulated domestic price of foreign capital goods, p_t^m, is 115 percent higher by 1991. This rise in input costs causes the relative price of investment, q_t, to increase (see equation 1.12). Our calibrated model delivers a 22.5 percent rise in the relative price of investment over this period. Moreover, annual output per effective unit of labor, y_t, is 2.9 percent lower in 1991 than in 1981. This loss in output arises from the cumulative impact of the rising relative scarcity of foreign capital goods which over time creates an increasing distortion in the production of investment goods. This

[14] An alternative assumption would have been to assume that the current policy remains in place forever, but this seems unreasonable since it would imply that the capital stock is driven to zero in the limit.

distortion is reflected in a steep rise in the relative price of investment. If government policy in India had merely increased the number of licenses to keep pace with the rise in productivity and workers, the model would have remained at the 1981 steady state and investment would not have become more expensive to produce over this period. The stability of the price of imported capital goods would have prevented the capital stock from lagging behind other inputs in the economy so that output per effective unit of labor would have been constant. The first column of Table 1.3 reports the change in the relative price of investment over the pre-reform period and compares them to the data.

Table 1.3: Change in q (%)

	Before reform	During reform
Data	44%	-26%
Model	23%	-28%

In order to show that this rise in q_t is not sensitive to the 1981 value of $\bar{M}_t$, we redo our quantitative analysis using $\bar{M}_t = 0.0220895$ which implies an initial $p^m = 2$ whereas $p^m = 1.72$ in the benchmark calibration.This change resulted in the model generating a rise in q_t of 22.57 percent instead of 22.54 percent.

1.4.2 During Reform

The Indian government instituted major reforms beginning in 1991 which included a quick dismantling of import controls on capital goods. Our goal is to focus on the impact of these specific reforms using our calibrated model. To implement this reform in our calibrated model, we assume that the restrictions on capital goods imports were fully in place during 1991 and completely removed by 1992. In addition to the removal of non-tariff barriers, the Indian government reduced tariff rates on capital goods from 72.7 percent in 1990 to 7.6 percent in 2006 (see Table 1.2). We assume for our quantitative analysis that all tariff changes are complete by 2006 and that there are no further changes so that a new steady state can be calculated at the lowest tariff rate in 2006. In order to extract the contribution of the removal of the quantity restrictions from the contribution of the tariff rate reductions to the relative price of investment in 1991, we calculate a hypothetical steady state for 1991. To obtain this steady state, we pick a tariff rate, θ^*, to match p^m to its pre-reform 1991 value, so that $p^{m^*} = p^w(1+\theta^*) = 3.71$, where the world price equals unity as usual. This hypothetical steady state corresponds to a situation where the actual tariff rate is equal to the implied distortion caused by the non-tariff barrier. One interpretation of this steady state is that the government has committed to keep the domestic price of foreign capital goods constant at the 1991 level by increasing licenses at the combined growth rate of productivity and employment. Having calculated key variables (y^*, c^*) in this steady state, we can compare their values to two other steady states, one in which tariff rates are reduced to 72 percent and another in which they are further reduced to the 2006 value of 7.6 percent. We can then calculate the total change in y and c by comparing the hypothetical 1991 steady state to the 2006 steady state and also the pure contribution of tariff rate reductions by comparing the steady state where tariff rates are fixed at the pre-reform weighted average of 72 percent to the 2006 steady state. Comparing the three steady states, we find that the total effect of the policy changes led to a fall in q_t of 28 percent. Out of this total fall, the tariff rate reduction alone accounts for 12 percent while 16 percent comes from the removal of quantity restrictions. The actual fall in the relative price of investment in India was 26 percent. The policy change induced fall in the domestic price of capital imports induces an increase in the import of capital goods used in investment goods creation, more capital accumulation and higher levels of output per unit of effective labor. In the new steady state, y is 17.8 percent higher than y^*. The pure contribution of the tariff rate reduction to this large rise in output per effective unit of labor is 6.5 percent while the remainder comes from the removal of quantity restrictions. Assuming no further declines in the tariff rate on capital goods imports, the policy change implies that consumption per worker is permanently higher by 13.4 percent compared to the hypothetical steady state in 1991.

1.4.2.1 Transition to the 2006 Steady State

While it is clear that the reduction in imported capital goods price lead to a permanently higher level of output and consumption, there could be welfare losses along the transition path as consumption falls in order to accumulate enough capital to reach the new steady state. In order to explore this issue and to further characterize the transition of the economy to the policy changes we calculate the transition path taking as given the 1991 capital stock, k_{1991}, obtained from the pre-reform transition. As observed in Table 1.2, tariff rate reductions occurred in a series of steps of unequal sizes. In order to deal with this, we compute the transition path during the reform period by assuming that tariff rates were reduced in three steps. These tariff rates for the transition are: i) 52 percent, ii) 23 percent, and iii) 7.6 percent.[15] The policy functions used during the reform period are calculated using the value function iteration (endogenous grid points method) proposed by Carroll (2006) (see Appendix 1.A for details). The transition follows the policy rule corresponding to a 52 percent tariff rate for 4 periods using k_{1991} as the starting value. Thereafter, the economy switches to a new policy rule corresponding to a 23 percent tariff rate for 9 periods using the 4th period capital stock as the starting value. After 9 periods the policy rule switches one last time to correspond to a 7.6 percent tariff rate using the existing capital stock in that period as the starting value. The economy follows this policy rule until it reaches the new steady state.

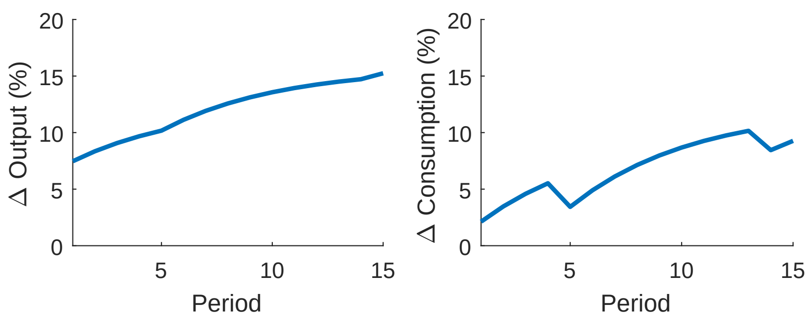

Figure 1.2: The impact of capital import reform in the benchmark economy

Figure 1.2 illustrates the first 15 periods of the transition paths for output per effective unit of labor, y_t and consumption per effective unit of labor, c_t expressed in percentage changes relative to the level of y^* and c^* respectively, in the hypothetical 1991 steady state discussed above. Output per effective unit of labor rises steadily with barely perceptible kinks around the shifts in policy rules. In contrast to this steady climb, c_t displays sharp kinks around shifts in policy rules but nonetheless rises relative to the 1991 steady state value. By 2006, y_t is 15.2 percent above and c_t is 9.3 percent respectively above the hypothetical steady state values but are quite far from reaching the 2006 steady state which takes over 60 periods to reach. Adding up the extra consumption in these periods, we find that the economy generates an additional 101 percent of steady state c^* over these 15 years. Similarly, over the 15

[15] The tariff rate falls from 72 percent to 52 percent between 1990 to 1992. It is on average 23 percent from 1993 to 2004. It is 7.6 percent in 2006. Data is not available for all years.

years, the economy produces an extra 180 percent of 1991 steady state output per effective unit of labor, y^*. Since output grows faster on the transition path than in the steady state (where all growth comes from exogenous sources) we can use the average growth rate of y_t during these 15 transition years to calculate the contribution of the reform to the observed increase in Indian GDP growth rate in the transition period. We find that y_t grows at 0.49 percent per annum on average in this period, which suggests that liberalization of capital goods imports may have, on its own, contributed one-fifth of the observed 2.2 percent rise in the growth rate of GDP per worker in India. This is interesting in light of the results in Bosworth and Collins (2008) that capital accumulation contributed 39 percent of the 4.6 percent growth in GDP per worker seen during the reform period in India.

1.4.3 Sensitivity Analysis

In this section we use a constant elasticity of substitution specification instead of a Cobb-Douglas specification for the investment good production technology. Table 1.4 shows the change in the relative price of investment induced by the calibrated model in both the pre-reform period and during the period of reform for different values of the CES parameter σ, expressed in elasticity of substitution. As σ is varied, we recalibrate the model to maintain the capital import to consumption ratio at the 1981 value for India which is 0.03. It shows that as the elasticity of substitution between domestic final goods and foreign capital increases, the responsiveness of the relative price of investment in the model decreases in both periods. The intuition is evident - if the investment goods producer can readily substitute foreign capital goods with domestic final goods, the impact of a rise in the price of imported capital can be mitigated. Given that the vast majority of capital goods are produced in just 10 nations (Mutreja et al., 2016), these substitution possibilities are likely to be quite limited in practice. As a result we view the benchmark Cobb-Douglas results, reproduced here, on the conservative end of the ability of our model to account for the relative price of investment movements in India.

Table 1.4: Sensitivity analysis: change in q (%)

	Data	Model		
		CES (Elasticity)		Cobb-Douglas
		0.80	1.33	
Before reform	44%	28%	16%	23%
During reform	-24%	-32%	-24%	-28%

1.5 Conclusion

In this paper, we construct a small open economy model where the government uses tariff and non-tariff barriers to limit the import of foreign capital goods. We calibrate the model to India using data from the Penn World Table and use it to account for the dramatic rise and fall of Indian relative price of investment. Our benchmark calibration implies that the model can generate a 23 percent rise in the relative price of investment between 1981 and 1991 due to increasing distortions created by quantitative restrictions on capital goods imports in the face of a growing economy. The model also accounts for a 28 percent fall in the relative price of investment over the subsequent 15 years as tariff rates fell from 72.7 percent to 7.6 percent and quantity restrictions were removed. The model allows us to separate the impact of tariff rate reductions from the impact of the implicit distortions to investment created by quantity restrictions on capital goods imports. We uncover a considerable general equilibrium impact of these price changes on output and consumption per worker and show that the Indian government's import substitution policies exerted a significant drag on the economy prior to reform. Moreover the removal of capital import restrictions and reduction of tariff rates accounts for one fifth of the observed increase in GDP per worker in India between 1991 and 2006.

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1.A Appendix

Algorithm for Calculating the Transition to the 2006 Steady State

We use value function iteration - endogenous grid points method (VFI-EGM) proposed by Carroll (2006) to solve the model. The recursive version of the problem is:

$$V(k) = \max_{c,k'} \log c + \beta V(k')$$

subject to the household's budget constraint, capital accumulation, all firms' first order conditions, and market clearing conditions. The following steps are used to solve the model and obtain transition to the 2006 steady state:

1. find policy rule corresponding to $\theta = 0.52$ using the following steps:
   - $p^m = p^w(1+\theta)$; p^w is normalized to 1,
   - calculate steady state, obtain k_{ss}, V_{ss}
   - create a fixed grid for $k' = [k'_{min}, k'_{max}] = [0.25 \times k_{ss}, 1.5 \times k_{ss}]$, where size of k' is $[100 \times 1]$
   - guess $\frac{\partial V(k')}{\partial k'} = V^k_{ss}$,
   - calculate q using

$$q = \frac{1}{\eta}\left(\frac{\eta p^m}{1-\eta}\right)^{1-\eta}$$

- calculate $\frac{d}{m}$ using

$$\frac{d}{m} = \frac{\eta}{1-\eta} p^m$$

- use the Newton method to solve for optimal k (i.e., k^{opt}) in following:

$$k^\alpha - \left(\frac{\beta}{\gamma_l \gamma_z q} E \frac{\partial V(k')}{\partial k'} \right)^{-1} - \left(\frac{d}{m} + p^w \right) \left(\frac{d}{m} \right)^{-\eta} (\gamma_l \gamma_z k' - (1-\delta)k) = 0$$

- calculate current period consumption, c^{opt} using

$$c^{opt} = \left(\frac{\beta}{\gamma_l \gamma_z q} E \frac{\partial V(k')}{\partial k'} \right)^{-1}$$

- get $\frac{\partial V(k^{opt})}{\partial k^{opt}}$ using envelop condition

$$\frac{\partial V(k^{opt})}{\partial k^{opt}} = \frac{1}{c^{opt}} (\alpha k^{opt(\alpha-1)} + q(1-\delta))$$

- interpolate k' to $\frac{\partial V(k')}{\partial k'}$ using the relationship between k^{opt} to $\frac{\partial V(k^{opt})}{\partial k^{opt}}$
- update the guess for $\frac{\partial V(k')}{\partial k'}$
- check convergence of $\frac{\partial V(k')}{\partial k'}$
- the policy rule is $[k', k^{opt}]$.

2. simulate y and c from period 1 to period 4 using k_{1991} as the starting value,

3. calculate the policy rule for using $\theta = 0.23$ following step 1,

4. simulate y and c for next 9 periods, i.e., from period 5 to 13 using the 4th period capital stock as the starting value,

5. calculate the policy rule for using $\theta = 0.076$ by following step 1,

6. simulate y and c next 87 periods using the 13th period capital stock as the starting value.

Chapter 2

Why is Agricultural Productivity So Low in Poor Countries? – The Case of India

2.1 Introduction

It is well known that low agricultural labor productivity is a major impediment to development. In India, labor productivity in agriculture is only a fifth of the level of urban productivity. In other words, agriculture seems to be one sector of the economy where ways of production lag particularly far behind the frontier technology, with small non-mechanized farms persisting through time (e.g. Foster and Rosenzweig (2011)). More generally speaking, the gap in agricultural labor productivity between the rich and poor countries is so large that it accounts for most of the observed income gap. Our main goal is to understand why poor countries fail to mechanize their ways of farming.

We begin with the premise that residing in a rural area provides access to a network that effectively insures its residents against income fluctuations. This premise has a solid foundation in a large body of literature and survey data.[1] If households indeed value their agricultural land beyond its productive value because it provides them with access to a network that effectively insures against labor income risk, then they are less willing to migrate to the city where labor earnings risk is uninsured. As a result, labor remains cheap in agriculture, and the incentives for switching to capital-intensive methods of farming stay weak. This description captures the main essence of our mechanism.

We capture this mechanism in a dynamic general equilibrium model that features uninsured labor income risk in the city. We allow for a very general production technology in agriculture that allows us to endogenize labor productivity through the choice of farm size and capital intensity. We assume that capital can substitute for labor, but land is a complementary input to both. There are diseconomies of scale in production, so reducing the number of farms in favor of larger farms raises productivity. As has become standard in literature on structural transformation, our framework features non-homothetic preference. Risk is completely insured away in rural areas. In urban areas, individuals self-insure against labor income risk through savings. Finally,

[1] Townsend (1994) has drawn attention to risk sharing feature in rural areas by empirically documenting that household consumption co-moves with village average consumption in India and shows little dependence on own income. Appendix 2.A.4 shows that risk sharing is much higher in rural areas compare to in urban areas in India. Santaeulàlia-Llopis and Zheng (2018) have found higher levels of consumption insurance in rural areas than in urban areas in China. De Magalhães and Santaeulàlia-Llopis (2018) have found more consumption insurance in rural areas than in urban areas in poor African countries.

newborns optimally choose whether to live in the rural or urban area.

The direct effect of the location choice is that an increase in the relative supply of urban labor will result in the fall of the urban-rural wage gap, mechanization of rural production, an increase in the average farm size and therefore labor productivity. The presence of non-insurable risk in urban labor income creates the urban-rural wage gap and consumption gap. In the scenario of no risk in urban areas, there is a spatial equilibrium with no consumption gap.

In order to understand the quantitative importance of our mechanism, we calibrate the model (under the assumption of a stationary steady state) to data for India for the period around 2000. We use wage data to discipline the labor market risk in the city. We rely heavily on the agriculture census of India to discipline the technology parameters. Our model successfully replicates the urban-rural wage gap (a factor of 1.3). To assess the importance of differential insurance access across locations, we study an abstract policy intervention. To be more precise, we employ the calibrated model to quantify the effect of introducing complete insurance in the city on migration and labor productivity in agriculture. As a result, the urban-rural consumption gap disappears and the share of workers in the agricultural area declines from 0.59 to 0.55, implying increased labor movement to urban areas. Our mechanism working through migration is consistent with Munshi and Rosenzweig (2016) which shows that informal insurance in rural areas decreases low-skilled labor migration in India. While this effect on labor reallocation is far from dramatic, the impact on agricultural productivity is very large. In the agricultural sector, capital input per farm rises by 120% as farm size expands and capital inflows to substitute for the lost labor. In fact, we find that the average farm size increases by 12 percent. The labor productivity gap between the two sectors decreases by 30 percent.

One clear implication of our benchmark model with uninsured labor risk in the city is the presence of the urban-rural consumption and wage gap. Because workers have differential access to social insurance across the two locations, they must be compensated with greater levels of consumption in the city. In India, the urban-rural wage gap declines from 1.7 to 1.3 and the consumption gap declines from 1.3 to 1.2 from 1983 to 2008 (Hnatkovska and Lahiri, 2016).[2] But these gaps are still large, and they cannot be explained by differences in observed worker characteristics or justified

[2]Wage gaps are obtained from a regression of (log) wages on a rural dummy, age, and age squared. Consumption gaps are obtained from a regression of (log) consumption expenditures on a rural dummy.

by worse amenities in the cities. Hnatkovska and Lahiri (2016) have examined the urban-rural both wage and consumption expenditure gaps in India by incorporating differential sectoral productivity shocks.Young (2013) has examined the urban-rural consumption expenditure gaps in 65 countries. Young (2013) and Hnatkovska and Lahiri (2016) have found the gaps cannot be accounted for by differences in observed characteristics of urban-rural workers, such as schooling levels. Gollin et al. (2017) have emphasized that spatial models require that amenities (or something else) must be worse in the urban areas to justify the presence of the observed urban-rural consumption inequality as an equilibrium outcome. Contrary to this implication, they show (using African data) that cities tend to offer not only higher consumption but also better quality amenities along all dimensions. Our framework based on differential access to insurance across the two locations provides one alternative explanation for the presence of the consumption gap that is consistent with spatial models. Another alternative explanation is of course the difference in unobserved worker characteristics across the two locations.

Our work is also related to Adamopoulos and Restuccia (2014), Adamopoulos and Restuccia (2018) and Chen et al. (2017). These papers focus on misallocation of inputs across farms of varying productivity. Adamopoulos and Restuccia (2014) have emphasized that resource misallocation across farms have the potential to account for differences in farm size and productivity between rich and poor countries. In contrast, we focus on trying to understand why the average farm size is small in India.

The rest of the paper is structured as follows. Section 1 presents the model while Section 2 discusses the calibration of model parameters. Section 3 presents quantitative results from the benchmark calibration as well as some sensitivity analysis. Concluding remarks are followed by an appendix that outlines model solution used in our paper.

2.2 The Model

Consider a model economy where time is discrete and indexed by $t = 0, 1, 2...$ N new households are born every period and live for exactly 2 periods (young and old). There are two spatially separated locations: rural and urban. We associate these locations with agricultural production and the "rest of the economy". Newborns decide once and for all on their location. The location determines the sector of employment and access to insurance. We denote by χ_t the fraction of generation t choosing to live in the rural area.

The timing of decision-making within each period is given as follows:

1. the new generation chooses their location of residence;

2. labor endowment shocks are realized;

3. in urban areas, identical CRS firms hire labor and rent capital to produce non-agricultural goods. In rural areas, households choose between running their own farms or working for wages, and farm managers hire labor and rent capital;

4. agricultural farms hire workers, rent land and capital to produce agricultural goods;

5. households receive proceeds for their factors of production and make consumption/savings decisions.

While we assume that people work where they live, we allow for capital to freely flow across locations. This means that rental rates of capital will equalize across locations, and this will be reflected in our notation from the start.

2.2.1 Urban Area

2.2.1.1 Urban Firms

The urban sector produces the non-agricultural good. It is comprised of a large number of identical firms endowed with a constant returns to scale technology. This allows us to restrict attention to a single aggregate firm that exhibits competitive behavior. The aggregate output of the non-agricultural good is given by $Y_{n,t} = A_n K_{n,t}^{\alpha} N_{n,t}^{1-\alpha}$, where $K_{n,t}$ and $N_{n,t}$ denote aggregate employment of capital and effective units of labor. A_n denotes total factor productivity. We set the non-agricultural good to be the numeraire so that all time t prices are quoted in the units of this good. Taking factor rental rates $w_{n,t}$ and the rental price of capital r_t as given, the aggregate firm hires inputs to maximize profit:

$$\max_{K_{n,t},\, N_{n,t}} \left\{ Y_{n,t} - w_{n,t} N_{n,t} - r_t K_{n,t} \right\}. \tag{2.1}$$

2.2.1.2 Urban Households

Households that choose to live in urban areas face idiosyncratic labor market risk. We follow the standard approach to model this risk as a stochastic endowment of effective labor units. The effective labor endowment when young and old are given by $\kappa\exp(\zeta^y)$ and $\kappa\exp(\zeta^o)$ where ζ^y is drawn from the set $[-\zeta,\zeta]$ with probabilities π^y and $1-\pi^y$, while ζ^o is drawn from the same set with probabilities depending on the previous realization. We assume the transition probability matrix of the form

$$\begin{bmatrix} \pi & 1-\pi \\ 1-\pi & \pi \end{bmatrix},$$

We also assume the initial probability distribution $[\pi^y, 1-\pi^y]$ is stationary with respect to the above transition probability matrix, i.e. $\pi^y = 0.5$. This approach of modeling labor endowment shocks allows us to approximate the AR(1) log earnings process estimated from the data. Because there is a continuum of young and old households, the probabilities also correspond to the overall measures of households experiencing a given shock. The measure of the young households with $\exp(-\zeta)$ endowment is always 0.5, and the measure of the old households with $\exp(-\zeta)$ is also 0.5 (due to the stationarity assumption). This immediately implies that the average labor endowment per young and per old household in each period is given by $\kappa\,(0.5)\,[\exp(-\zeta)+\exp(\zeta)]$ and we normalize $\kappa = \frac{1}{(0.5)[\exp(-\zeta)+\exp(\zeta)]}$ as to make the average labor endowment equal to 1.

Expected utility, prior to the initial shock realization, of the cohort born in t choosing the urban location, is given by

$$\begin{aligned} EU_n \;=\; & E_{\zeta^y}\Bigg\{\phi\frac{(a^y_{n,t}(\zeta^y)-\bar{a})^{1-\sigma}}{1-\sigma}+(1-\phi)\frac{c^y_{n,t}(\zeta^y)^{1-\sigma}}{1-\sigma}+ \\ & \beta E_{\zeta^o|\zeta^y}\left(\phi\frac{(a^o_{n,t+1}(\zeta^y,\zeta^o)-\bar{a})^{1-\sigma}}{1-\sigma}+(1-\phi)\frac{c^o_{n,t+1}(\zeta^y,\zeta^o)^{1-\sigma}}{1-\sigma}\right)\Bigg\}, \end{aligned} \tag{2.2}$$

where $a^y_{n,t}(\zeta^y)$ and $c^y_{n,t}(\zeta^y)$ denote the initial state-contingent consumption of the agricultural and non-agricultural goods when young, $a^o_{n,t+1}(\zeta^y,\zeta^o)$ and $c^o_{n,t+1}(\zeta^y,\zeta^o)$ denote the state history-contingent consumption of agricultural and non-agricultural good when old, ϕ denotes the preference weight on agricultural goods, and β is the discount factor. We follow the large literature on structural transformation in assuming non-homothetic preferences stemming from the subsistence consumption level $\bar{a} > 0$.

We now describe the state-contingent budget constraints. The household is born with no capital. There is no leisure in utility, so the household inelastically supplies the available endowment of effective labor units to the urban firm. In period t, the young agent earns wage income $w_{n,t}\exp(\zeta^y)$ and allocates it between consumption and savings, $k^n_{t+1}(\zeta^y)$. In period $t+1$, the old agent earns labor and capital income, all of which he consumes at this point. The following constraints must hold for all t and all possible realization histories (ζ^y, ζ^o):

$$p_{a,t}a^y_{n,t}(\zeta^y) + c^y_{n,t}(\zeta^y) + k^n_{t+1}(\zeta^y) = w_{n,t}\kappa\exp(\zeta^y), \tag{2.3}$$

$$p_{a,t+1}a^o_{n,t+1}(\zeta^y,\zeta^o) + c^o_{n,t+1}(\zeta^y,\zeta^o) = w_{n,t+1}\kappa\exp(\zeta^o) + r_{t+1}k^n_{t+1}(\zeta^y), \tag{2.4}$$

where $p_{a,t}$ denotes the price of agricultural goods in time t. Note we assume capital fully depreciates upon use which is reasonable given that the model period length is half the length of one's working years.

Note there are no social insurance arrangements in the city. Households can self-insure against labor market risk, which gives rise to precautionary savings. This assumption is consistent with the lack of social insurance institutions typically observed in developing countries.

2.2.2 Agricultural Area

One defining characteristic of rural residence that we aim to capture is that it provides access to a network of friends and family that effectively insures against idiosyncratic labor income risk. The straightforward way to model this is to introduce idiosyncratic labor endowment risk as in the urban households' problem described above, but allow for households to enter risk-sharing contracts which are perfectly enforceable. If we assume that households are born identical, this would yield identical allocations for all households regardless of idiosyncratic realization. To simplify the exposition, we go directly to modeling a representative household that faces no uncertainty. Each household has 1 unit of time endowment when young and when old. We assume that young households work for wages, while old households either work for wages or manage farms.

Land is in fixed supply denoted by L. It is initially in the hands of the initial old residents of the agricultural areas. The old households sell land to the young at the end of the period.

2.2.2.1 Agricultural Farms

Suppose an endogenous fraction ε_t of old households use their entire time endowment to manage farms while the remaining households work for wages. Given the rental price of capital, land and labor $(r_t, q_t, w_{a,t})$, each manager hires capital, labor and land $(k^f_{a,t}, n^f_{a,t}\, l^f_{a,t})$ to maximize profit, denoted by d_t:

$$\max_{k^f_{a,t}, n^f_{a,t}, l^f_{a,t}} d_t = p_{a,t} y_{a,t} - w_{a,t} n^f_{a,t} - r_t k^f_{a,t} - q_t l^f_t\,, \tag{2.5}$$

where $y_{a,t}$ is the amount of non-agricultural good produced. We modify the Lucas (1978) span of control approach to endogenize the farm size and assume

$$y_{a,t} = A_a \left[(1-\theta) \left(l^f_t \right)^{\rho} + \theta \left(\nu \left(k^f_{a,t} \right)^{\mu} + (1-\nu) \left(n^f_{a,t} \right)^{\mu} \right)^{\frac{\rho}{\mu}} \right]^{\frac{\eta}{\rho}},$$

where $\eta \in (0,1)$ represents the span of managerial control parameter, μ governs the elasticity of substitution between capital and labor, and ρ governing the elasticity of substitution between land and the capital-labor composite, $\theta \in (0,1)$ captures the relative importance of land and $\nu \in (0,1)$ determines the relative importance of capital and labor. To be consistent with well-documented empirical facts regarding the elasticity of substitution between factors of production in agriculture in developing countries, we assume $\mu > 0$ and $\rho < 0$, i.e. capital and labor are gross substitutes whereas land is a complementary factor (e.g. Salhofer (2000)).

We require that managers are indifferent between working for wages and running farms, which yields the no-arbitrage condition:

$$d_t = w_{a,t}. \tag{2.6}$$

2.2.2.2 Agricultural Households

Households that choose to live in agricultural areas have the same preferences as households residing in the urban area, except they face no uncertainty. Again, this is a shortcut to modeling fully enforceable risk-sharing contracts. The lifetime utility of a cohort born in t that chose the agricultural location is given by

$$U_a = \phi\,\frac{(a^y_{a,t}-\bar{a})^{1-\sigma}}{1-\sigma} + (1-\phi)\,\frac{\left(c^y_{a,t}\right)^{1-\sigma}}{1-\sigma} + \beta\left\{\phi\frac{(a^o_{a,t+1}-\bar{a})^{1-\sigma}}{1-\sigma} + (1-\phi)\frac{\left(c^o_{a,t+1}\right)^{1-\sigma}}{1-\sigma}\right\}, \tag{2.7}$$

where $a^y_{a,t}$ denotes consumption of the agricultural good at young age in period t, $c^y_{a,t}$ denotes consumption of the non-agricultural good at young age in period t, $a^o_{a,t+1}$ denotes consumption of the agricultural good at old age in period $t+1$, $c^o_{a,t+1}$ is consumption of the non-agricultural good at old age in period $t+1$.

The household is born with no capital and no land. Therefore, the only source of income for the young agents is labor income. The young purchase agricultural and non-agricultural goods for consumption and save in the form of capital and land. They purchase $k_{a,t+1}$ units of capital from the non-agricultural good producer at price 1 and l_{t+1} units of land from the old households at price $p_{l,t}$. The time t budget constraint for the young agents is summarized as follows:

$$p_{a,t}a^y_{a,t} + c^y_{a,t} + p_{l,t}l_{t+1} + k_{a,t+1} = w_{a,t}. \tag{2.8}$$

In period $t+1$, the old agents either work for wages or manage firms, in either case earning $w_{a,t+1}$. They also get rental income from capital $r_{a,t+1}k_{a,t+1}$, rental income from land $q_{t+1}l_{t+1}$ and income from land sale $p_{l,t+1}l_{t+1}$. The old agents consume all of their income. The time $t+1$ budget constraint of the old agents is given by

$$p_{a,t+1}a^o_{a,t+1} + c^o_{a,t+1} = w_{a,t+1} + r_{a,t+1}\,k_{a,t+1} + q_{t+1}l_{t+1} + p_{l,t+1}l_{t+1}. \tag{2.9}$$

2.2.3 Equilibrium

Measure χ_t of each cohort decides to locate in agricultural areas. This measure is determined by equalization of lifetime utility across areas:

$$EU_n = U_a. \tag{2.10}$$

Before we can define our market clearing conditions, we need to establish additional notation. Recall that N denotes the size of each cohort. The measure of young households living in urban locations is then given by $N^y_{n,t} = \chi_t N$. Because all young households survive to adulthood and the location choice is permanent, this is also the measure of the urban old agents in time $t+1$:

$$N_{n,t}^{y} = N_{n,t+1}^{o} = \chi_t N.$$

Likewise, the measure of the rural young and old households in time t and $t+1$ is given by

$$N_{a,t}^{y} = N_{a,t+1}^{o} = (1-\chi_t)\, N.$$

The measure of farm managers in time t is given by $\varepsilon_t N_{a,t}^{o}$.

Given the nature of our study, it suffices to focus on a stationary equilibrium, i.e. a decentralized competitive equilibrium characterized by stationary allocations, prices and location choice.

Definition A stationary equilibrium is defined as state-contingent allocations for the urban young households $\{a_n^y(\zeta^y),\, c_n^y(\zeta^y),\, k_n(\zeta^y)\}_{\zeta^y}$ and for the urban old households $\{a_n^o(\zeta^y,\zeta^o), c_n^o(\zeta^y,\zeta^o)\}_{(\zeta^y,\zeta^o)}$, for the rural area households $\{a_a^y,\, c_a^y,\, k_a,\, l,\, a_a^o, a_a^o\}$, for the urban firm $\{Y_n, K_n, N_n\}$ and for the rural farms $\{y_a, k_a^f,\, n_a^f,\, l^f, d\}$, prices $\{w_n,\, w_a,\, r,\, q,\, p_l\ ,\, p_a\}$, the measure of each cohort choosing to live in the urban area (χ) and the measure of the rural young managing farms (ε) such that:

1. given the equilibrium prices, the allocations for the urban households maximize utility (2.2) subject to the budget constraints (2.3) and (2.4);

2. given the equilibrium prices, the allocations for the rural households maximize utility (2.7) subject to the budget constraints (2.8) and (2.9);

3. given the equilibrium prices, the allocation for the urban firm maximizes profit given in (2.1);

4. given the equilibrium prices, the allocation for the rural firm maximizes profit given in (2.5);

5. lifetime utility equalizes across the two locations: equation (2.10) holds;

6. farm managers are indifferent between managing a farm and working for wages, i.e. (2.6) holds;

7. all markets clear:

- labor market in agriculture:

$$\varepsilon N_a^o n_a^f = N_a^y + N_a^o(1-\varepsilon), \tag{2.11}$$

- labor market in the urban area:

$$N_n = N_n^y + N_n^o, \tag{2.12}$$

- capital market:

$$K_n + \varepsilon N_a^o k_a^f = N_n^o k_n + N_a^o k_a, \tag{2.13}$$

- land market in agriculture:

$$\varepsilon N_a^o l^f = N_a^o l = L, \tag{2.14}$$

- agricultural goods market:

$$\varepsilon N_a^o y_a = N_n^y a_n^y + N_n^o a_n^o + N_a^y a_a^y + N_a^o a_a^o, \tag{2.15}$$

- non-agricultural goods market:

$$Y_n = N_n^y c_n^y + N_n^o c_n^o + N_n^y k_n + N_a^y c_a^y + N_a^o c_a^o + N_a^y k_a, \tag{2.16}$$

where $N_n^y = N_n^o = \chi N$ denote the population size of young and old residing in the urban area, and $N_a^y = N_a^o = (1-\chi)\, N$ denote the population size of young and old residing in the agricultural area, and where $k_n = 0.5\left[k_n\left(\zeta^y = -\zeta\right) + k_n\left(\zeta^y = \zeta\right)\right]$ is investment per urban young household (and capital holdings per urban old household), $a_n^y = 0.5\left[a_n^y\left(-\zeta\right) + a_n^y\left(\zeta\right)\right]$ is the average consumption of the agricultural good for the urban young household, $c_n^y = 0.5\left[c_n^y\left(-\zeta\right) + c_n^y\left(\zeta\right)\right]$ is the average consumption of the non-agricultural good for the urban young household, $a_n^o = 0.5[\pi a_n^o\left(-\zeta,-\zeta\right) + \left(1-\pi\right) a_n^o\left(-\zeta,\zeta\right) + \pi a_n^o\left(\zeta,\zeta\right) + \left(1-\pi\right) a_n^o\left(\zeta,-\zeta\right)]$ is the average consumption of the agricultural good for the urban old household, and $c_n^o = 0.5[\pi c_n^o\left(-\zeta,-\zeta\right) + \left(1-\pi\right) c_n^o\left(-\zeta,\zeta\right) + \pi c_n^o\left(\zeta,\zeta\right) + \left(1-\pi\right) c_n^o\left(\zeta,-\zeta\right)]$ is the average consumption of the non-agricultural good for the urban old household.

A few clarifications are in order. Note that the demand for labor appearing on the left hand side of (2.11) comes from the number of individual farms ($=\varepsilon N_a^o$), each of which demands n_a^f units of labor. While the young supply their labor inelastically, only a measure $1-\varepsilon$ of the old households work for wages – the rest manage farms. The demand for capital on the left hand side of equation (2.13) comes from the aggregate firm in the urban area (K_n) and measure εN_a^o of individual farms each demanding k_a^f. The supply of capital is from urban and agricultural households. Taking expectations over shock realizations gives the average savings by the young. These are last period savings of the current old. Just like the capital market, there is a single market for each of the consumption goods. Expectations are used to get the appropriate averages. It should be noted that capital investment comprises a part of the demand for the non-agricultural good.

The characterization of equilibrium is provided in Appendix 2.A.1. Here we focus on the key conditions of the model. Performing the optimization yields to the following first-order conditions describing households intratemporal and intertemporal trade-offs in urban area,

$$\begin{aligned}
\frac{\phi}{1-\phi}\left(\frac{c_n^y(\zeta^y)}{a_n^y(\zeta^y)-\bar{a}}\right)^{\sigma} &= p_a, \text{ for } \zeta^y = -\zeta, \zeta \\
\frac{\phi}{1-\phi}\left(\frac{c_n^o(\zeta^y,\zeta^o)}{a_n^o(\zeta^y,\zeta^o)-\bar{a}}\right)^{\sigma} &= p_a, \text{ for } (\zeta^y,\zeta^o) = (-\zeta,-\zeta),(-\zeta,\zeta),(\zeta,\zeta),(\zeta,-\zeta)
\end{aligned}$$

$$\begin{aligned}
\frac{c_n^y(-\zeta)^{-\sigma}}{\pi c_n^o(-\zeta,-\zeta)^{-\sigma} + (1-\pi)\, c_n^o(-\zeta,\zeta)^{-\sigma}} &= \beta r, \\
\frac{c_n^y(\zeta)^{-\sigma}}{\pi c_n^o(\zeta,\zeta)^{-\sigma} + (1-\pi)\, c_n^o(\zeta,-\zeta)^{-\sigma}} &= \beta r,
\end{aligned}$$

and in agricultural area,

$$\begin{aligned}
\frac{\phi}{1-\phi}\left(\frac{c_a^y}{a_a^y-\bar{a}}\right)^{\sigma} &= p_a \\
\frac{\phi}{1-\phi}\left(\frac{c_a^o}{a_a^o-\bar{a}}\right)^{\sigma} &= p_a
\end{aligned}$$

$$\left(\frac{c_a^y}{c_a^o}\right)^{-\sigma} = \beta r.$$

If we eliminate of risk in urban areas, i.e. by setting $\zeta = 0$ and $\pi = 1$, intratemporal and intertemporal tradeoffs become:

$$\frac{c_n^y(0)}{a_n^y(0) - \bar{a}} = \frac{c_n^o(0,0)}{a_a^o(0,0) - \bar{a}} \quad = \quad \frac{c_a^y}{a_a^y - \bar{a}} = \frac{c_a^o}{a_a^o - \bar{a}},$$

$$\frac{c_n^y(0)}{c_n^o(0,0)} = \frac{c_a^y}{c_a^o}.$$

So elimination of risk in urban areas eliminates consumption gaps across urban and agricultural areas.

In agricultural area, the optimal measure of manager (ε) depends on

$$d = w_a,$$

which implies managers must be indifferent between managing a farm and working. Then for a given value of ε, the average farm size (l_f) is determined by following land market clearing condition:

$$l_f = \frac{L}{\varepsilon N_o^a}.$$

Since the average farm size is a ratio of land to measure of manager, the elimination of idiosyncratic risk in urban areas will drive a lower employment in agricultural area and also lower the measure of manager, that will increase the average farm size.

2.3 Calibration

Our objective is to investigate quantitatively the extend of an abstract policy intervention – a provision of complete insurance against earnings risk in urban area. The general strategy is to calibrate the benchmark model parameters by assuming steady state of our model adequately represents India's economic scenario during the period of 2000–2012, and by targeting several moments during this period.

Our strategy is to conduct the proposed quantitative analysis by choosing some parameters based on a priori information, by finding the rest as a part of solution in the benchmark model. We set values for ζ and π to match the estimated wage premium in non-agricultural sector and the estimation steps are detailed in Appendix 2.A.3. In agricultural production technology, we use the value for ρ is -2.0 to maintain the elasticity of substitution of land with composite of capital and labor below 0.5.[3] We set $\theta = 0.5$ assuming equal the relative importance of land and capital-labor composite. We set $\mu = 0.6$ which governs the elasticity of substitution between capital and labor, $\nu = 0.5$ assuming equal the relative importance of capital and labor. We also set $\eta = 0.8$ assuming average profit is 20 percent of agricultural output. For preference, we use $\sigma = 2$ and for time discounting factor $\beta = 0.422$.[4] Maintaining values for ζ, π, μ, ν, θ, η, σ, β, we choose parameters (α, $\bar{a}$, A_n, L, N) to match the following empirical moments: $[i] : \frac{rK_n}{Y_n} = 0.33$; $[ii] : 1 - \chi = 0.59$; $[iii] : w_n/w_a = 1.35$; $[iv] : \frac{p_a(a_a^y + a_a^o)}{c_a^y + c_a^o + p_a(a_a^y + a_a^o)} = 0.50$; and $[v] : l_f = 1.33$. These moments refer to, in the order listed, capital income share in non-agriculture sector, population share in rural area, urban-rural wage gap, consumption expenditure share of agricultural good in rural area, and the average farm size. The resulting calibrated parameter values are in Table 2.1 and the data source of the targeted moments are detailed in the Appendix 2.A.2. We normalize the value for A_a is 1. Table 2.2 shows the remaining parameters values used in the benchmark model.

Table 2.1: Calibrated parameters

α	$\bar{a}$	A_n	L	N
0.33	0.001	2.0202	8.4296	20

Table 2.2: Remaining parameters

ϕ	β	σ	ρ	μ	ν	θ	η	A_a	ζ	π
0.4	0.422	2	-2	0.6	0.5	0.5	0.8	1	0.6596	0.5961

In Table 2.3, column 2 shows data moments, and column 3 shows model produced moments. This shows our model successfully match empirical moments. The model also matches well urban-rural consumption gap which is not targeted in calibration.

[3]Salhofer (2000) reports that the empirical estimates of the elasticity of substitution in agriculture generally fall well below 0.5.

[4]Household's optimal decision in agricultural area is $\frac{c_a^y}{c_a^o} = \beta r$. Assuming in long run $\frac{c_a^y}{c_a^o} = 1$ and risk free annual interest rate is four percent, then $\beta = \frac{1}{1.04^{22}}$ with each period is 22 years long.

Table 2.3: Model: moments

	(1)	(2) Data	(3) Benchmark Model	(4) Counter-factual
Targeted Moment				
[*i*] Capital income share	$\frac{rK_n}{Y_n}$	0.33	0.33	0.33
[*ii*] Employment share	$1-\chi$	0.59	0.59	0.55
[*ii*] wage gap	w_n/w_a	1.35	1.46	1.00
[*iv*] Consumption expenditure share of agricultural good in rural area	$\frac{p_a(a_a^y+a_a^o)}{c_a^y+c_a^o+p_a(a_a^y+a_a^o)}$	0.50	0.48	0.52
[*v*] Average farm size (Hectares)	l_f	1.33	1.34	1.50
Non-Targeted Moment				
[*vi*] Consumption gap	$\frac{c_n^y+c_n^o+p_a(a_n^y+a_n^o)}{c_a^y+c_a^o+p_a(a_a^y+a_a^o)}$	1.25	1.56	1.00

Notes: The data sources of above moments are detailed in Appendix 2.A.2.

2.4 Result

To examine the quantitative effect of an abstract policy intervention, we shut down idiosyncratic labor endowment risk in the non-agricultural area by setting $\zeta = 0$ and $\pi = 1$. Then we solve the model by maintaining the same values for the remaining parameters. In Table 2.3, column 4 shows model generated moments of this counterfactual exercise. Before explaining result, we define some key variables: Aggregate Agricultural Output, $Y_a = \varepsilon(1-\chi)Ny_a$; Aggregate Labor in the Agricultural Area, $N_a = (1-\chi)2N$; Aggregate Capita in the Agricultural Sector, $K_a = \varepsilon(1-\chi)Nk_a^f$, real GDP $= Y_n + p_aY_a$, Labor Productivity Gap $= \frac{Y_n/N_n}{p_aYa/N_a}$; Labor Productivity in the Agricultural Sector $= \frac{p_aY_a}{N_a}$; Aggregate Productivity $= \frac{Y_n+p_aY_a}{2N}$. Due to full insurance, the consumption are equalized between two sectors and the wage premium disappears in urban areas, i.e. wages in two sectors are equalized. The population share in agricultural area decreases from 59 percent to 55 percent which means workers relocate to the urban area. The demand for capital per farm increases by 120 percent, whereas the demand for labor per farm increases by 3.6 percent due to substitutability between capital and labor inputs in agricultural good production technology. The average farm size increases by 12 percent. Capital flows to the urban sector so the ratio K_n/K_a decreases by 47 percent. Labor productivity in agricultural sector increases by 37 percent and aggregate labor productivity increases by 16 percent, the labor productivity gap between two sectors falls by 30 percent and GDP rises by 17 percent. Table 2.4 summarizes main results.

Table 2.4: Model: main results

	w_a	$\frac{Y_n/N_n}{p_aYa/N_a}$	$\frac{p_aY_a}{N_a}$	$\frac{Y_n+p_aY_a}{2N}$	$\frac{K_n}{K_a}$	GDP	k_a	l_f
% change in counter-factual relative to benchmark model	42	-30	37	16	-47	17	120	12

Notes: $Y_a = \varepsilon(1-\chi)Ny_a$, $N_a = (1-\chi)2N$, $K_a = \varepsilon(1-\chi)Nk_a^f$, GDP $= Y_n + p_aY_a$, Productivity Gap $= \frac{Y_n/N_n}{p_aYa/N_a}$, Productivity in agricultural sector $= \frac{p_aY_a}{N_a}$, Aggregate Productivity $= \frac{Y_n+p_aY_a}{2N}$.

To summarize: when a complete insurance is implemented in the urban area, capital mobilizes from rural to urban area, labor moves from rural to urban area, it generates higher labor productivity in agricultural sector through raising farm size. Given its parsimonious nature, we deem this quantitative result to be a considerable success of the model, and help to inquire further importance of social insurance policy in the city.

2.5 Conclusion

There is a large productivity gap between urban and agricultural sectors in India. Furthermore, agricultural production is characterized by small non-mechanized farms. We develop a tractable quantitative framework by incorporating one potential explanation. If residing in a village provides access to a network that effectively insures against income fluctuations, then households are less willing to live in the cities where labor income risk is uninsured. As a result, labor stays cheap in agriculture, and the incentives for switching to capital-intensive methods of farming remain weak. In order to understand the quantitative importance of this mechanism, we calibrate the model to Indian data and study an abstract policy intervention – a provision of complete insurance against earnings risk in the city. The policy intervention reduces the urban-rural labor productivity gap by 30 percent and raises aggregate labor productivity by 16 percent. This effect comes about because of the 7 percent drop in agricultural share of employment, which encourages an inflow of capital in the agricultural sector and raises the average farm size by 12 percent.

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2.A Appendix

2.A.1 Model Solution

The unknowns are:

– $\{a_n^y(-\zeta),\, c_n^y(-\zeta),\, k_n(-\zeta),\, a_n^y(\zeta),\, c_n^y(\zeta),\, k_n(\zeta)\}$ (6 variables)

– $\{a_n^o\left(-\zeta,-\zeta\right), a_n^o\left(-\zeta,\zeta\right), a_n^o\left(\zeta,-\zeta\right), a_n^o\left(\zeta,\zeta\right), c_n^o\left(-\zeta,-\zeta\right), c_n^o\left(-\zeta,\zeta\right), c_n^o\left(\zeta,-\zeta\right), c_n^o\left(\zeta,\zeta\right)\}$ (8 variables)

– $\{a_a^y,\, c_a^y,\, k_a,\, l\}$ (4 variables)

– $\{a_a^o, a_a^o\}$ (2 variables)

– $\{Y_n,\, K_n,\, N_n\}$ (3 variables)

– $\{y_a,\, k_a^f,\, n_a^f,\, l^f,\, d\}$ (5 variables)

– $\{w_n,\, w_a,\, r,\, q,\, p_l\ ,\, p_a\}$ (6 variables)

– χ , ε (2 variables)

So we have 36 unknowns.

2.A.1.1 Urban Firm

Given r, w_n, we solve for $\{Y_n, K_n, N_n\}$.The aggregate urban firm profit maximization yields the following 3 conditions:

$$\begin{aligned} Y_n &= A_n K_n^{\alpha} N_n^{1-\alpha}, \\ w_n &= (1-\alpha) A_n K_n^{\alpha} N_n^{-\alpha}, \\ r &= \alpha A_n K_n^{\alpha-1} N_n^{1-\alpha}. \end{aligned}$$

2.A.1.2 Individual Farms

Given r, w_a, p_a, we solve for $\{y_a, k_a^f,\, n_a^f,\, l^f,\, d\}$. Individual farm profit maximization yields the following 5 conditions:

$$\begin{aligned} y_a &= A_a \left[(1-\theta) \left(l^f\right)^{\rho} + \theta \left(\nu \left(k_a^f\right)^{\mu} + (1-\nu) \left(n_a^f\right)^{\mu} \right)^{\frac{\rho}{\mu}} \right]^{\frac{\eta}{\rho}}, \\ w_a &= \theta p_a \eta A_a \left[\cdot\right]^{\frac{\eta}{\rho}-1} \left(\nu \left(k_a^f\right)^{\mu} + (1-\nu) \left(n_a^f\right)^{\mu} \right)^{\frac{\rho}{\mu}-1} (1-\nu) \left(n_a^f\right)^{\mu-1}, \\ r &= \theta p_a \eta A_a \left[\cdot\right]^{\frac{\eta}{\rho}-1} \left(\nu \left(k_a^f\right)^{\mu} + (1-\nu) \left(n_a^f\right)^{\mu} \right)^{\frac{\rho}{\mu}-1} \nu \left(k_a^f\right)^{\mu-1}, \\ q &= p_a \eta A_a \left[\cdot\right]^{\frac{\eta}{\rho}-1} (1-\theta) \left(l^f\right)^{\rho-1}, \\ d &= (1-\eta) p_a y_a. \end{aligned}$$

2.A.1.3 Urban Area Households

Given p_a, r, w_n, the urban household variables, $\{a_n^y(-\zeta),\, c_n^y(-\zeta),\, k_n(-\zeta),\, a_n^y(\zeta),\, c_n^y(\zeta),\, k_n(\zeta)\}$ and $\{a_n^o\left(-\zeta,-\zeta\right),\; a_n^o\left(-\zeta,\zeta\right),\; a_n^o\left(\zeta,-\zeta\right),\; a_n^o\left(\zeta,\zeta\right),\; c_n^o\left(-\zeta,-\zeta\right),\; c_n^o\left(-\zeta,\zeta\right),\; c_n^o\left(\zeta,-\zeta\right),\; c_n^o\left(\zeta,\zeta\right)\}$ (14 variables), must satisfy the following conditions:

– 6 budget constraints:

$$\begin{aligned} p_a a_n^y(\zeta^y) + c_n^y(\zeta^y) + k_n(\zeta^y) &= w_n \kappa \exp(\zeta^y), \text{ for } \zeta^y = -\zeta, \zeta \\ p_a a_n^o(\zeta^y, \zeta^o) + c_n^o(\zeta^y, \zeta^o) &= w_n \kappa \exp(\zeta^o) + r_n k_n(\zeta^y), \text{ for } (\zeta^y, \zeta^o) = \{(-\zeta, -\zeta), (-\zeta, \zeta), (\zeta, \end{aligned}$$

– 6 conditions involving the marginal rate of substitution between c and a at each tree node:

$$\begin{aligned} \frac{\phi}{1-\phi}\left(\frac{c_n^y(\zeta^y)}{a_n^y(\zeta^y) - \bar{a}}\right)^\sigma &= p_a, \text{ for } \zeta^y = -\zeta, \zeta \\ \frac{\phi}{1-\phi}\left(\frac{c_n^o(\zeta^y, \zeta^o)}{a_n^o(\zeta^y, \zeta^o) - \bar{a}}\right)^\sigma &= p_a, \text{ for } (\zeta^y, \zeta^o) = (-\zeta, -\zeta), (-\zeta, \zeta), (\zeta, \zeta), (\zeta, -\zeta) \end{aligned}$$

– 2 conditions involving the intertemporal expected rate of substitution between c^y and c^o:

$$\begin{aligned} \frac{c_n^y(-\zeta)^{-\sigma}}{\pi c_n^o(-\zeta, -\zeta)^{-\sigma} + (1-\pi)\, c_n^o(-\zeta, \zeta)^{-\sigma}} &= \beta r \text{ for } \zeta^y = -\zeta \\ \frac{c_n^y(\zeta)^{-\sigma}}{\pi c_n^o(\zeta, \zeta)^{-\sigma} + (1-\pi)\, c_n^o(\zeta, -\zeta)^{-\sigma}} &= \beta r \text{ for } \zeta^y = \zeta \end{aligned}$$

2.A.1.4 Rural Area Households

Given p_a, r, w_a, q, p_l, the rural area household variables, $\{a_a^y,\ c_a^y,\ k_a,\ l\}$ and $\{a_a^o,\ a_a^o\}$ (6 variables), must satisfy the following conditions:

– 2 budget constraints:

$$\begin{aligned} p_a a_a^y + c_a^y + p_l l + k_a &= w_a, \\ p_a a_a^o + c_a^o &= w_a + r\, k_a + (q + p_l)\, l. \end{aligned}$$

– 2 conditions involving the marginal rate of substitution between c and a at each age:

$$\begin{aligned} \frac{\phi}{1-\phi}\left(\frac{c_a^y}{a_a^y - \bar{a}}\right)^\sigma &= p_a \\ \frac{\phi}{1-\phi}\left(\frac{c_a^o}{a_a^o - \bar{a}}\right)^\sigma &= p_a \end{aligned}$$

– 1 condition involving the intertemporal rate of substitution between c^y and c^o :

$$\left(\frac{c_a^y}{c_a^o}\right)^{-\sigma} = \beta r$$

– Note the household will save in terms of capital only if $r > \frac{q+p_l}{p_l}$, and in terms of land only otherwise. We will focus on the interior solution, which means that the rates of return equalize across the two assets, i.e.

$$r = \frac{q+p_l}{p_l},$$

we can solve the above equations to determine a_a^y, c_a^y, a_a^o, a_a^o and savings $s_a = p_l l + k_a$.

2.A.1.5 Market Clearing

To summarize what we have so far, for given prices, $\{w_n,\ w_a,\ r,\ q,\ p_l\ ,\ p_a\}$, we have solved for all household and firm variables (except l and k_a) and we have obtained 2 additional conditions:

$$\begin{aligned} r &= \frac{q+p_l}{p_l}, \\ s_a &= p_l l + k_a, \end{aligned}$$

where s_a is known.

It remains to find $\{w_n,\ w_a,\ r,\ q,\ p_l,\ p_a\}$, l, k_a, χ, ε (10 variables).

The remaining conditions are 6 market clearing where we substituted for the population size of the young and old in urban areas, $N_n^y = N_n^o = \chi N$, and for the population size of young and old in agricultural areas $N_a^y = N_a^o = (1-\chi) N$:

– Labor market in agriculture:

$$\varepsilon (1-\chi) N n_a^f = (1-\chi) N (2-\varepsilon).$$

– Labor market in the urban area:

$$N_n = \chi 2N.$$

– Capital market:

$$K_n + \varepsilon (1 - \chi) N k_a^f = \chi N k_n + (1 - \chi) N k_a.$$

– Land market in agriculture:

$$\varepsilon (1 - \chi) N l^f = (1 - \chi) N l = L.$$

– Agricultural goods market:

$$\varepsilon (1 - \chi) N y_a = \chi N \left(a_n^y + a_n^o\right) + (1 - \chi) N \left(a_a^y + a_a^o\right).$$

– Non-agricultural goods market:

$$Y_n = \chi N \left(c_n^y + c_n^o + k_n\right) + (1 - \chi) N \left(c_a^y + c_a^o + k_a\right).$$

$k_n = 0.5 \left[k_n \left(\zeta^y = -\zeta\right) + k_n \left(\zeta^y = \zeta\right)\right]$ is investment per urban young household (and capital holdings per urban old household),

$a_n^y = 0.5 \left[a_n^y (-\zeta) + a_n^y (\zeta)\right]$ is the average consumption of the agricultural good for the urban young household,

$c_n^y = 0.5 \left[c_n^y (-\zeta) + c_n^y (\zeta)\right]$ is the average consumption of the non-agricultural good for the urban young household,

$a_n^o = 0.5[\pi a_n^o (-\zeta, -\zeta) + (1 - \pi) a_n^o (-\zeta, \zeta) + \pi a_n^o (\zeta, \zeta) + (1 - \pi) a_n^o (\zeta, -\zeta)]$ is the average consumption of the agricultural good for the urban old household,

$c_n^o = 0.5[\pi c_n^o (-\zeta, -\zeta) + (1 - \pi) c_n^o (-\zeta, \zeta) + \pi c_n^o (\zeta, \zeta) + (1 - \pi) c_n^o (\zeta, -\zeta)]$ is the average consumption of the non-agricultural good for the urban old household.

Measure χ of each cohort decides to locate in agricultural areas. This measure is determined by equalization of lifetime utility across areas:

$$EU_n = U_a.$$

We also know

$$d = w.$$

2.A.2 Data Source

[*i*] *Capital Income Share in Non-agriculture Sector*: Standard in sectoral model.

[*ii*] *Population Share in Rural Area*: Figure 1.6 of State of Indian Agriculture 2015-16 reports workforce engaged in agricultural sector is 0.59.

[*iii*] *Urban-rural Wage Gap*: We estimate the value for rural-urban wage gap by using Indian Human Development Survey I and II. Estimation is detailed in Appendix 2.A.3.

[*iv*] *Consumption Expenditure Share of Agricultural Good in Rural Area*: Anand and Prasad (2010) estimated minimum consumption requirement value to be 50 percent of food consumption for a sample of six emerging economies, including India.

[*v*] *The Average Farm Size*: Table 9.4 of State of Indian Agriculture 2015-16 reports the averages farm size is 1.33 Hectares,

[*vi*] *Consumption Gap per Capita*: Figure 2 in Hnatkovska and Lahiri (2016) reports the urban-rural consumption gap is 1.25 in 1999-2000.

2.A.3 Wage Process

We use data from Indian Human Development Survey (IHDS), conducted by University of Maryland and the National Council of Applied Economic Research, which is a nationally representative multi-topic panel survey. The first round (IHDS-I) was completed in 2004-05 and the second round (IHDS-II) was completed in 2011-12. We estimate the persistence of wage for employed male worker living in urban area.

- Wage samples include only male, ages from 16 to 65, male household head, not enrolled in educational institution, full time employed. We convert hourly wage to daily wage in IHDS data.

- We estimate following equation for workers in urban area:

$$\log w_{i,t} = \beta_1 + \beta_2 \text{age}_{i,t} + \beta_3 \text{age}_{i,t}^2 + \varepsilon_{i,t}^w,$$

- then calculate residuals $r^w_{i,t}$:

$$r^w_{i,t} = \log w_{i,t} - \log \bar{w}_t,$$

- then estimate following equation, get ρ and calculate σ^2 (variance of $\varepsilon_{i,t}$)[5] ,

$$r^w_{i,t} = \alpha + \rho r^w_{i,t-1} + \varepsilon_{i,t}.$$

- Using estimated $\rho = 0.575^3$ and $\sigma^2 = 0.419$, we calculate ζ and π by following:

$$\zeta = \sqrt{\frac{\sigma^2}{1-\rho^2}},$$

$$\pi = \frac{1+\rho}{2}.$$

- We estimate wage gaps ($\exp(\beta_4) = 1.37$) in year 2012:

$$\log w_i = \beta_1 + \beta_2 \text{age}_i + \beta_3 \text{age}_i^2 + \beta_4 \text{sector} + \varepsilon^w_i.$$

2.A.4 Measuring Consumption Insurance

We use consumption and income data from IHDS-I and IHDS-II, consumption and income are measured as follows:

- **Consumption**: The adult-equivalent consumption measures is computed by dividing the household consumption measure by the equivalence scales (KP) in Krueger and Perri (2006), defined as follows:

$$KP = [(\text{\# of adults age} \geq 15) + 0.7 \times (\text{\# of children age} < 15)]^{0.7}.$$

- **Income**: The adult-equivalent household income is obtained by dividing the benchmark income by the number of working age adults (age ≥ 15)

We estimate a partial insurance model by following steps[6]:

[5]We need to use three exponents of ρ to get model equivalent.

[6]Estimation steps are similar to Santaeulàlia-Llopis and Zheng (2018), Storesletten et al. (2004).

1. regress the (logged) adult-equivalent income Y_t on dummies of age, education level, province of residence, and ethnic minority separately by rural/urban status and by year,

$$Y_{i,t} = \beta_1 + \beta_2 age_{i,t} + \beta_4 education_{i,t} + \beta_4 state_{i,t} + \beta_5 minority_{i,t} + \varepsilon^Y_{i,t},$$

2. regress the (logged) adult-equivalent consumption measure C_t on dummies of age, education level, province of residence, and ethnic minority separately by rural/urban status and by year,

$$C_{i,t} = \beta_1 + \beta_2 age_{i,t} + \beta_4 education_{i,t} + \beta_4 state_{i,t} + \beta_5 minority_{i,t} + \varepsilon^C_{i,t},$$

3. estimate residual of income y_t and consumption c_t ,

$$y_{i,t} = Y_{i,t} - \bar{Y}_{i,t},$$

$$c_{i,t} = C_{i,t} - \bar{C}_{i,t},$$

4. take the difference in the residuals; unexplained income growth Δy_t and unexplained consumption growth Δc_t ,

$$\Delta y_{i,t} = y_{i,t} - y_{i,t-1},$$

$$\Delta c_{i,t} = c_{i,t} - c_{i,t-1}$$

5. regress Δc_t on Δy_t, estimate shock transmission parameter ψ and estimate variance of the residuals (see Table 2.5).

Table 2.5: Regression results, IHDS, 2005, 2012

Dependent variable: Δc		
	Rural	Urban
Independent variable: Δy		
ψ	0.102	0.194
	(0.008)	(0.011)
adjusted R^2	0.028	0.075
N	13540	5755

Chapter 3

Baby Bonus, Anyone? Examining Heterogeneous Responses to a Pro-Natalist Policy

3.1 Introduction

With declining birth rates in most of the developed world, nations are concerned with the burden placed on the working population to support a growing fraction of the retired population. Understanding the potential problems of below-replacement fertility rates raises a number of questions about pro-natalist policies: do they work, who is taking advantage of the incentives provided, and how costly are the programs? Past research finds that tax exemptions on children, child tax credits, and family allowances all increase fertility; however large increases in these benefits would be needed to reach replacement fertility levels (Zhang et al., 1994). Evidently, the policies are very expensive and if we can observe heterogeneous responses to these incentives then governments could tailor pro-natalist policies to encourage births from certain groups at lower cost.

The Canadian province of Quebec implemented a universal cash transfer, namely, the Allowance for Newborn Children (ANC), for all babies born from May 1988 to September 1997 to all residents of Quebec. Hereafter, we refer to this transfer as a baby bonus. This quasi-natural experiment has many qualities that allow us to estimate the impact of financial incentives on fertility.[1] First, the structure and payment plan of the pro-natalist policy was announced suddenly in the newspapers, allowing it to be treated as an unanticipated exogenous shock to the people of Quebec (La Presse, 1988; Montreal Gazette, 1988a,b).[2] The front page of the Montreal Gazette read "Have more babies, Liberals say". Second, the baby bonus reached as high as C$8,000 for families having a third child or higher. This is a sizable benefit and not tied to any other benefits or clawed back at higher income brackets. Parent and Wang (2007) stress the importance of fiscal incentives being large enough to induce an increase in household births. Also, our control group, the province of Ontario did not introduce new child benefit legislation until 1997, allowing for a clean comparison (Battle and Mendelson, 1997; Milligan and Stabile, 2011). Finally, and most importantly, the ANC is a universal pro-natalist policy implemented specifically in response to low fertility rates. Many baby bonuses are implemented for specific subgroups of the population,

[1] There are many papers studying the impact of fiscal incentives on fertility; examples include Ang (2015), Baughman and Dickert-Conlin (2009), Brewer et al. (2012) , Cohen et al. (2013), González (2013), LaLumia et al. (2015), and Raute (2017).

[2] Unfortunately, the cancellation of the policy is announced well in advance, and replaced by universal childcare; this creates a less credible experimental environment at the end of the policy period.

usually low-income individuals, to promote horizontal equity. For example, the Canada Child Tax Benefit payment, an in-cash transfer for anyone with a child, is reduced once adjusted family net income is over a threshold income (Milligan, 2016b). Since the ANC is universal we are able to examine the heterogeneous response of different subgroups of the population to this pro-natalist policy, and, thus, shed light on which women are being induced to have more children, and how family formation is being shaped.

Assuming pro-natalist policies do impact fertility, it is important to know whether or not the effect is permanent or transitory. If the effect is transitory this implies that women only choose to adjust the timing of their births, while this could impart a shift on the age distribution, if the government is trying to increase family size then resources are being wasted on a purely transitory effect. If the effect is permanent this implies that women did choose to have more babies, and, hence, increases completed fertility. Past papers are unable to answer whether this baby bonus had a permanent effect because enough time needs to pass to examine the entire child-bearing period of each cohort. Using the confidential birth vital statistics and census data, we are able to calculate completed fertility rates for a number of cohorts that were impacted by the ANC, thereby providing the first analysis on whether or not the ANC had a permanent or transitory effect on fertility. The ANC has been previously studied by Duclos et al. (2001), Milligan (2005), Kim (2012, 2014), Ang (2015). The latter studies build on Duclos et al. by using an additional data set, which contains demographic information about the mother and the family. All papers find a positive average effect of the ANC on fertility, Duclos et al. using vital statistics, while Milligan and Kim are using the public-use census file to control for individual household characteristics. Unfortunately, the public-use census files contain a small sample and indicate only if a child under the age of six is present on the census day, not the actual age. With access to de-identified individual census records we know the exact date of birth. Moreover, unlike past papers, a larger sample size, in addition to detailed data, allows us to examine the heterogeneous response to the ANC by parity (birth order), sibship sex composition, income, education, and immigrant status and to estimate meaningful marginal effects. Ang (2015) addresses the effect of the ANC on birth order using the confidential census file, but does not delve into the spacing of births, changes in completed fertility, or family formation as we do and does not delve into the heterogenous effect on different groups except for birth parity.

Since we know the sex of older siblings in the household, we ask whether the sex of the two older children influences the decision to have a third, something suggested by the large literature on parental preference. To our knowledge, this is the first paper to examine the effect of cash transfers on sibship sex composition. There is a primary preference for one-of-each-sex with a secondary preference for a son in North America and other developed countries as opposed to a strong primary preference for sons in developing countries (e.g., Andersson et al. (2006); Angrist and Evans (1998); Freedman et al. (1960); Ost and Dziadula (2016); Williamson (1983)). We find evidence that the baby bonus is able to alleviate some of these sex preferences through the large cash incentive for higher parity births. Specifically, we find that a third birth is more likely when there are two previous sons or a previous son and daughter than if both are daughters. Also, we find that parents having a previous son are more likely to have another child with the baby bonus comparing to having a previous daughter. That is consistent with studies in both Canada and the United States that find a gender preference for sons exists (Almond et al., 2013; Dahl and Moretti, 2008). Our results remain the same under various specifications and sensitivity tests. We also find that the baby bonus produced more three-child households with one daughter and two sons.

In addition to finding heterogeneous responses to the ANC by parity, and sibship sex composition, we also find a hump shape response by income group, and a positive response by maternal education. We also confirm, both graphically and through regression analysis, that the baby bonus created both a transitory and permanent effect; Quebec women not only chose to have their children sooner, but also to have more children. Thus, the increase in completed fertility rates implies that the ANC was successful in its endeavor to increase fertility.

The next section of the paper discusses theoretical considerations, while section 3.3 explains the institutional background of the ANC. Section 3.4 and 3.5 examine the two datasets and the empirical methods respectively. In section 3.6.2 we discuss our results, followed by a conclusion in section 3.7.

3.2 Theoretical Considerations

According to Becker (1960), policy changes that increase incomes, reduce the price of an additional child, or both, would be expected to increase fertility. However, that view was revised: such policy changes may not lead to an increase in the number of children if there is a meaningful trade-off between child quantity and quality (Becker, 1981; Becker and Lewis, 1973). Furthermore, a price change would alter quality unless quantity and quality are strong complements in parental utility functions. Consequently, theoretical considerations lead to ambiguous predictions of fertility responses to reforms. This also illustrates why heterogeneous responses to a pro-natal policy are inevitable and need to be included in theoretical models.[3] The traditional quantity-quality trade-off model is proving less clear-cut for the developed world, and not empirically evident (Angrist et al. (2010)). By way of example, highly educated women do not necessarily plan to have fewer children than their less educated counterparts (Esping-Andersen, 2009; Kravdal and Rindfuss, 2008).

Cash transfers may not encourage more births if parents already reach or exceed their optimal number, as may be common in developing countries (Palermo et al., 2016). However, if they would like to have more children, as may be more common in developed countries, cash transfers might have positive effects on fertility. Some research finds that highly educated women's desired number of children is greater than their actual number. For example, Testa (2014) finds a positive association between women's level of education and lifetime fertility intentions at both the individual and country levels. While highly educated people intend to have more children than less educated women (Heiland et al., 2008) they ultimately have fewer (Bongaarts, 2001; Quesnel-Vallée and Morgan, 2003). Such findings are consistent with a negative relationship between maternal education and fertility and imply that the marginal effect of an incentive such as a baby bonus may be higher for more educated women. This is not to say that highly educated women would have more children than those less educated, but their marginal effect in response to a baby bonus may be higher. Consistent with recent literature, we also find that more highly educated women respond to the ANC more than those less educated.[4]

According to Cigno and Ermisch (1989) a rise in child benefits would increase completed fertility, but the tempo of fertility and the amount spent on each child would fall. While the empirical finds little evidence on an impact on completed fertility

[3] New theoretical models are accounting for observed heterogeneous effects. For example, to account for the effect of a child-care policy on fertility, Yakita (2018) allows for responses to differ by level of maternal education.

[4] Shang and Weinberg (2013) study the case in the United States. Raute (2017) finds that an earnings-dependent maternity leave benefit in Germany increases fertility most among the middle and upper end of the education and income distributions.

many papers find that tempo effects rise. That is consistent with Parent and Wang's (2007) model of fertility decisions with liquidity constraints: the child benefit must be substantial to induce a rise in completed fertility. As Cigno and Ermisch (1989) note, if the assumption of access to the capital markets does not apply, as in the case of young couples, child benefits would be expected to raise the tempo of childbearing. Interestingly, we find both an increase in quantum and tempo effects from the baby bonus, suggesting that the cash incentive we analyse was strong enough to increase completed fertility as well as shorten the time between births.

A discrepancy between theory and empirics can be caused by differing designs of child benefits and how family policies are constructed is of great importance. Many create "speed premiums" which essentially encourage women to space their births closer together in order to take advantage of a benefit (see Björklund, 2006; Lalive and Zweimüller, 2009; Neyer and Andersson, 2008). Since women in developed countries have a strong preference for two children (Berrington, 2004), the Quebec government tailored its baby bonus to encourage fertility by offering more generous transfers at higher parity births. Thus, we find large differential effects by parity, specifically for third and higher births. By contrast, Cygan-Rehm's (2016) finding of no differential birth-order response to a German reform is not surprising, given that the payments were the same across parities. Naturally, the timing and number of births differ for women with different levels of education and family income, due to differing opportunity costs and thus differing marginal prices for children.

As argued in the Becker and Lewis (1973) seminal paper, parents trade-off the number of children they have with the quality of those children. If families trade-off quality for quantity (Mogstad and Wiswall, 2016; Pop-Eleches, 2006) and low-income families are sensitive to these pro-natalist policies then these baby bonuses may worsen intergeneration inequality. That is, if low-income parents are induced to have more children through pro-natalist policies then the quantity-quality trade-off suggests that these parents invest less in their children. Building on the quantity-quality theory, Becker and Tomes (1976) outline a U-shape model for the desired number of children as a function of income. This means that at low income levels, the overall income elasticity of demand for children is negative, whereas at high income levels it is positive. Their model predicts that an exogenous shock reducing the price of children would have low-income mothers spend extra income on children they already have rather than having more children because the substitution and the income effects work in opposite directions. This suggests that a baby bonus may not induce low-income families to have more children. On the other hand, the fixed baby bonus may not translate into a large enough percentage increase in income to induce high-income families to have another child. Thus, we expect the marginal effect of the baby bonus to decline at the upper end of the income distribution. In alignment with the theory, we find a hump shape response to the ANC by income group: there is little response among low-income families, mid-income families respond the most, and high-income families respond the least. This result contrasts with Milligan's (2005) finding of an overall positive response to income.[5]

Recent work by Riphahn and Wiynck (2017) examines the 1996 German child benefit program and finds that there is no fertility effect for low income couples. Further evidence from the UK found no increase in births among single women when a reform targeted at low-income households was implemented in 1999 (Brewer et al. (2012)). Also Moffitt (1994) and Hoynes (1995) find that the US Aid to Families with Dependent Children (AFDC) benefits had no impact on fertility for single mothers. It appears that the low price response among low income women may be because they spend any additional income on the children they already have rather than increasing the size of the family. Milligan (2005) also comments that more educated women may have more 'planned' pregnancies, and so are more responsive to price signals. Finally, in Cigno's (1986) theoretical model with endogenous fertility, if the wages of husbands and wives are positively correlated and families are differentiated only by earning ability, child benefits do not lead to greater inequality. If earning ability is highly correlated with education level, our result matches this case where higher child benefits do not increase the number of low income children.

3.3 Institutional Background

The ANC was a non-taxable in-cash transfer to all legal residents of Quebec that had a newborn, or adopted a child under the age of five, between May 1, 1988 and September 30, 1997. The amount of the benefit depended on the parity (birth order)

[5] Milligan (2005) estimates a probit regression with the variable "family income". He finds an overall positive coefficient, whereas we subgroup family income and estimate the same model to find the marginal effects of each income subgroup. Here, we are able to find a hump shape response for family income.

of the child.[6] The amount and exact timing of these payments are in Table 3.1. Also, the value of the benefit for third or higher order children continuously rose over the policy period. By the end of the policy parents of three or more children received C$8000, which, according to Milligan (2005), accounts for around 30 percent of the direct cost of the first five years of a child's life. Not surprisingly, the policy became expensive to continue, costing over C$1.4 billion between 1989 and 1996 according to Milligan (2002).[7] In September 1997, with the termination of the universal ANC, the provincial government instead implemented a universal C$5 a day childcare policy to encourage mothers' participation in the labor force. Also, the ANC was replaced with a new means-tested family allowance focusing on low-income families (Milligan and Stabile, 2011).

Table 3.1: Benefit payments under the allowance for newborn children

	First child	Second child	Third or higher child
May 1988 to April 1989	C$500 at birth	C$500 at birth	8 quarterly payments of C$375=C$3000
May 1989 to April 1990	C$500 at birth	C$500 at birth, C$500 on 1st birthday	12 quarterly payments of C$375=C$4500
May 1990 to April 1991	C$500 at birth	C$500 at birth, C$500 on 1st birthday	16 quarterly payments of C$375=C$6000
May 1991 to April 1992	C$500 at birth	C$500 at birth, C$500 on 1st birthday	20 quarterly payments of C$375=C$7500
May 1992 to Sept. 1997	C$500 at birth	C$500 at birth, C$500 on 1st birthday	20 quarterly payments of C$400=C$8000

Notes: Each cell reports the payments made for a child born within the specified period. Source: Milligan (2005).

Using the Canadian Tax and Credit Simulator (Milligan, 2016a), we calculate the total family benefits across different birth parities for different income levels in Quebec and Ontario from 1985 to 2000.[8] We observe total family benefits in the first year a child is born, across different birth parities, in Quebec and Ontario for a family income

[6]The baby bonus was paid to all births that were registered; we find no evidence of differences in ANC take-up rates by income.

[7]We confirm this calculation.

[8]The total family benefits include all refundable credits from federal government and provincial government. See figures 3.4 and 3.4 for the comparison of family benefits between Ontario and Quebec families.

of C\$20,000 and C\$60,000, respectively.[9] Both figures show that total family benefits are significantly higher in Quebec than Ontario during the baby bonus, and largest for third or higher birth parities during the ANC policy period. Furthermore, during the sample period, Ontario did not have any provincial baby bonus policies enacted.[10]

During the almost decade-long duration of the ANC policy two other policies could have potentially affected the number of births in Quebec. First, abortions were decriminalized in Canada following the strike down of Section 251 by the Supreme Court in 1991 with regards to *R. v. Morgentaler* (1988). The fading stigma of abortions can potentially influence fertility; however, the rate of abortion per 100 live births in Quebec showed only a slight increase between 1986 and 1992, from 14.7 to 16.6.[11] Moreover, we check to ensure that there is a parallel trend and that our difference-in-differences model is not contaminated by varying abortion rates between Quebec and Ontario. Second, Quebec was given constitutional power with regards to immigration in the Canada-Quebec Accord of 1991 (Young, 1998). If there is a difference in the fertility behavior of immigrants selected by Quebec, then variation from immigrants' fertility is misleadingly assumed to be attributed to the ANC instead. We address this concern and find that the exclusion of immigrants results in the same findings; we conclude that the Canada-Quebec Accord of 1991 does not affect our analysis. We also examine the response of the ANC by immigrant status and find that immigrant and non-immigrant families respond similarly.

3.4 Datasets

In this section we first describe the birth vital statistics dataset and discuss our graphical findings. Then we describe the census dataset that we use for regression analysis in Section 3.4.2.

[9] In simulation, if applicable, we assume that the second child is 6 years old and the third child is 10 years old.

[10] In 1997, after our sample period, Ontario introduced a means-tested child care supplement for working parents (Milligan and Stabile, 2011).

[11] Source: Statistics Canada. Table 106-9013

3.4.1 Vital Statistics

Using the confidential birth vital statistics from 1974 to 2011 we know the province of each birth, the mother's age, and the parity.[12] With these three critical variables we are able to look at trends in fertility between Quebec and Ontario to assess the impact of the ANC.[13] Figure 1 shows the total fertility rate (TFR), the cross-section of the sum of age-specific fertility rates in each year from 1974 to 2011, for women between the ages of 15 and 49.[14] The figure makes a very compelling argument for the positive effect the ANC had in Quebec.[15] The TFR in Quebec diverges from Ontario in the early 1980s, remains for 5 years, and then displays a narrowing of this gap starting in 1988.[16] Since the exact structure and payment plan of the baby bonus was not announced until the provincial budget speech of May 1988, the slight increase of births in 1988 could not have been affected by the ANC policy. However, Ontario also illustrated an increase in the same year, albeit not as steep as in Quebec. Furthermore, in the previous year's budget speech, the Quebec Minister of Finance, Gérard D. Levesque, announced that family assistance was an important aspect of the new budget, with a specific mention of allowances for families with three or more children being considered (Bernard (1989)). Perhaps, some families may have anticipated that a baby bonus of some sort would be implemented shortly.

Although the termination of the ANC is not experimentally ideal due to the introduction of universal childcare and the change in Ontario's child benefit policy,

[12]With the confidential data we are able to look at annual TFR for each year of age, whereas past papers using the public use data have had to use five-year age intervals.

[13]From all the Canadian provinces, the province of Ontario is the most comparable to Quebec; they are neighbors, as well as the two most populated provinces in Canada. There are many cities and towns on the border of these two provinces, and in one instance they even share the same metropolitan area (Ottawa-Gatineau).

[14]See Hotz et al. (1997) for a detailed comparison on total fertility rates (TFR) and completed fertility rates (CFR).

[15]In addition to graphical findings, we estimate a difference-in-differences (DID) model using the TFR as the outcome of interest for Quebec and Ontario with 1995 as the treatment year and 1988 as the comparison year; The DID model results in a 0.11 increase in the number of children born to Quebecois women in the treated year. As Manski and Pepper (2018) point out, such DID estimates require strong assumption on DID invariance. Following Manski and Pepper, we apply a class of the bounded-variation assumptions. We use the data prior to 1988 to calculate the bound parameter of bounded time variation, bounded inter-province variation, and bounded DID variation. The bounded DID estimates are between 0.104 and 0.199. These models are available upon request.

[16]We also compared Quebec to the Rest of Canada and find that it closely mirrors that of Ontario illustrating that the gaps we are observing in Quebec are not just in comparison to Ontario.

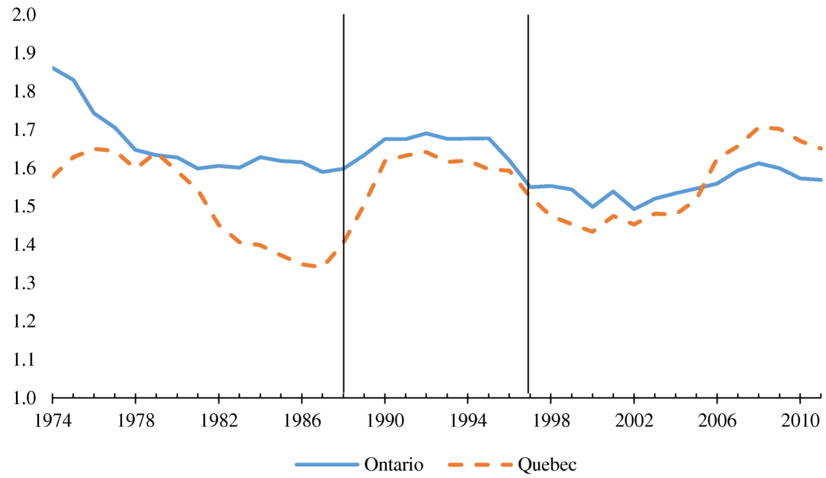

Figure 3.1: Total fertility rate, age 15–49. The first vertical bar signifies the start of the ANC policy in May 1988, and the second vertical bar signifies the end of the policy in September 1997. Source: Birth Vital Statistics, 1974 to 2011

we do see some evidence that Quebec's TFR fell immediately after its cancellation.[17] Figure 2 further decomposes the TFR by birth order. Here we observe Quebec first-order births surpass Ontario during the policy period. We also suggest that first-order births respond immediately to the policy, followed by second, and then third and higher. This illustrates parents having more children during the policy window in order to receive the substantially higher baby bonus for third and higher children.[18]

Many studies examining family policies usually find a transitory (tempo) effect (see Björklund, 2006; Cygan-Rehm, 2016; Heckman and Walker, 1990). It seems far easier to influence when a woman will have a child as opposed to how many. Parent and Wang (2007) examine Quebec after a family allowance that took place in Canada in the 1970's. Here they find only a transitory effect and specifically no quantum effect. They stress that the price change induced by the reform may simply not have been strong enough to cause a permanent effect. More specifically, with regards to this baby bonus, past work has shown that a transitory effect exists, however we are the first to

[17]Although the universal childcare policy is announced to start at the same time the baby bonus is canceled, no new subsidized childcare spaces were created before 2001 (Haeck et al. (2015)). In Norway, Havnes and Mogstad (2011) find that formal childcare acts as a substitute for informal childcare (arrangements with relatives, friends, and so forth) instead of encouraging new female labor force participation. Baker et al. (2008) examine childcare use in Quebec and do find some crowding out of existing arrangements is evident.

[18]Milligan (2002) writes that the rate for third and subsequent births in Quebec increased by 35 percent, from 0.217 per woman in 1987 to 0.294 in 1993, while falling elsewhere in Canada by 3 percent.

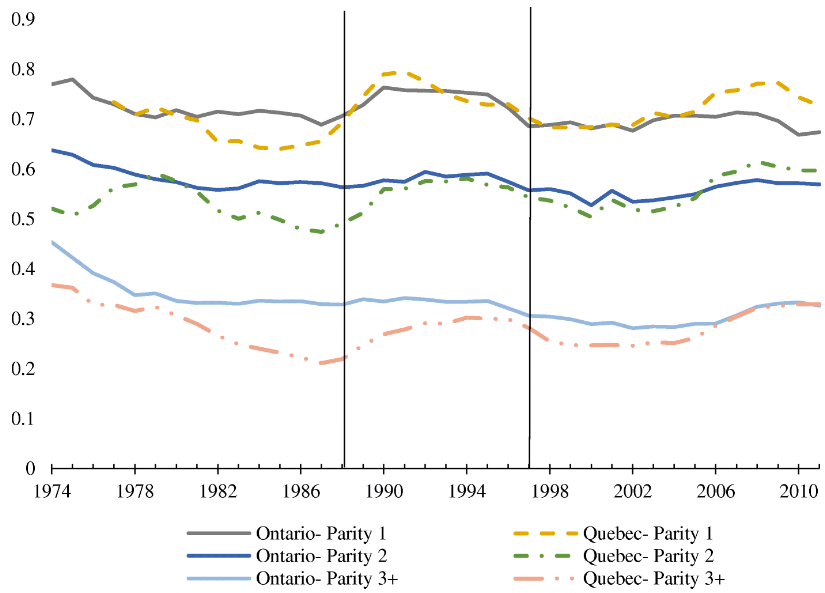

Figure 3.2: Total fertility rate by birth order, age 15–49. The first vertical bar signifies the start of the ANC policy in May 1988, and the second vertical bar signifies the end of the policy in September 1997. Source: Birth Vital Statistics, 1974 to 2011

explore whether this policy had a permanent effect on fertility as well.[19]

To assess the permanent impact of the ANC policy we need to look to the completed fertility rates. Figure 3 displays the CFR's for both Quebec (solid line) and Ontario (dashed line) starting from cohorts born in 1935. For the cohorts born from the late 1930's to the late 1950's the Ontario and Quebec completed fertility rates run in parallel, with Quebec lower by about 0.35 children per woman. These cohorts were either not affected by the policy, or were in the latter end of their childbearing years. For later cohorts, the ones that would have been most affected by the ANC, the gap narrows and then disappears altogether for the cohort born in 1970. For younger cohorts, born after 1970, we actually see Quebec's CFR surpass Ontario's by 3.5 percent. Prior to the policy, Quebec's CFR was on a steady decline dipping to around 1.58, after the baby bonus we see Quebec's slope is no longer parallel with Ontario and instead witness its CFR climb to 1.75 indicating a permanent effect of the policy

[19] We also show a transitory effect took place both graphically and in regression analysis. Our graphical results can be seen in the Appendix, Figure A3. These figures show the birth cumulative distribution function for each of three cohorts by age of mother and parity, separately for Ontario and Quebec. The difference between Ontario and Quebec is most evident for the third child, where one can observe the "middle cohort" in Quebec having children much earlier than their Ontario counterparts.

on their completed fertility. Quebec's CFR may well rise higher since there are more cohorts that have been exposed to the baby bonus that cannot presently be calculated.

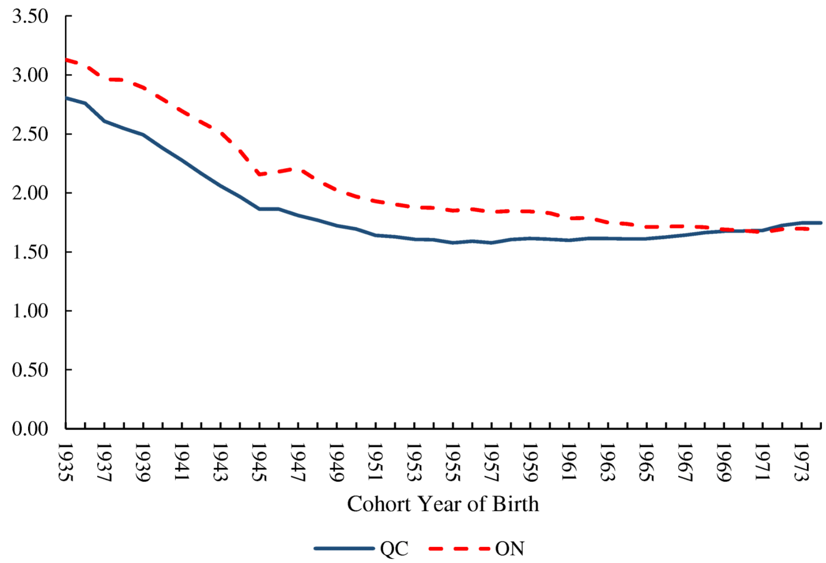

Figure 3.3: Completed fertility rate, cohorts aged 15–39. Birth Vital Statistics, 1950 to 2013

3.4.2 Census Data

The Canadian Population Census is conducted every five years; it provides household information recorded on Census Day. Our main results are based on the 1991 Census and the 1996 Census. We also use the 1986 Census for sensitivity analysis and the 2001 Census to analyze completed fertility rates. With the de identified files we observe the exact year a child is born.[20] To create a control group we choose to examine all married or common-law women from 1987 and 1988 from the 1991 census file. Since the policy was announced in the spring of 1988, any mothers who would have been incentivized by the ANC would have given birth, at the earliest, nine months later, which falls into 1989. Therefore, the closest control group to the start of the ANC is all married or common-law women in 1987 and 1988. When conducting robustness

[20]The main shortcoming of Milligan's (2005) study of the ANC is that the public-use census does not provide year of birth. This meant that the ANC policy period overlaps the 1991 census window, making it difficult to disentangle which births are part of the policy period. Also, the public-use census file has a very small sample size and does not allow for a thorough examination of heterogeneous effects like the confidential census file.

checks we change our control group to examine, for example, the number of births that take place in 1984 and 1985 from the 1986 Census. Our treatment group is all married or common-law women from 1994 and 1995 that live in Quebec. We use only two years for the treatment group so that the time period is balanced with our control group and for three additional reasons.[21] First, after seven years of the policy, every family should be familiar with the ANC, and would have had time to exploit it should they wish to. That is, the choice of treatment period avoids a heterogeneous information problem (e.g., more educated households know about the policy earlier than others). Second, it is before the cancellation of the policy was announced, thus residents are unaware of a possible change to the policy. Third, by choosing 1994 and 1995 from the 1996 Census, our income and household characteristics are from 1995, allowing for the most accurate estimates.

We limit our sample to married or common-law females between the age of 15 and 34, who have not changed provinces in the last five years prior to Census day, who are residents of Canada, and have positive income as defined below.[22] Only 11 percent of births from Quebec women under the age of 35, who have not changed provinces in the last five years, were to single women. We remove single mothers for two reasons: (1) we are trying to create a homogeneous group of women to compare, and (2) due to the way we define income all single mothers' income calculation would be misleading in our model.[23][24] We define income to be equal to the spouse's wage and self-employment income plus all investment income from both the spouse and woman.[25] [26] Female wages are excluded because of the endogeneity between female labor force participation and fertility decisions. In addition to income, we observe birth order and whether or not the family lives in an urban versus rural neighborhood

[21] As a robustness check we also use a three year and five-year window. See Section 6, B. Sensitivity Analysis for more detail.

[22] We limit the sample to women aged 34 to ensure we can identify all children; there is a concern that if the woman is older than 34 years of age she may have children living outside the home.

[23] As a robustness check, we examine the effect of the baby bonus on all single women. See Section 6, B. Sensitivity Analysis for more detail.

[24] Another reason we only look at married women is because we do not want to model the relationship between the decision to be married and fertility as studied in Baudin et al. (2015).

[25] We use the Canadian Consumer Price Index (CPI) for each province to convert nominal income into real income in 1992 constant Canadian dollars.

[26] The approach of using husband's income to measure family income has been adopted by many in the literature (See Hotz et al. (1988); Milligan (2005); Jones and Tertilt (2008)).

Table 3.2: Census summary statistics

	Quebec			Ontario		
	1986	1991	1996	1986	1991	1996
Had a child	0.239	0.207	0.268	0.269	0.236	0.282
Zero older children	0.497	0.633	0.488	0.46	0.603	0.471
One older child	0.232	0.194	0.243	0.225	0.193	0.234
Two or more older children	0.272	0.174	0.27	0.316	0.205	0.296
Female: 15–24 years old	0.207	0.176	0.16	0.192	0.136	0.112
Female: 25–29 years old	0.390	0.369	0.333	0.379	0.378	0.333
Female: 30–34 years old	0.404	0.455	0.507	0.43	0.487	0.556
Female: allophone	0.051	0.053	0.074	0.093	0.143	0.169
Female: francophone	0.869	0.88	0.862	0.031	0.056	0.052
Female: anglophone	0.081	0.063	0.065	0.877	0.799	0.78
Female: high school dropout	0.259	0.211	0.156	0.277	0.198	0.148
Female: high school diploma	0.231	0.193	0.166	0.216	0.216	0.17
Female: some post-secondary	0.42	0.479	0.504	0.386	0.445	0.494
Female: university degree	0.089	0.118	0.176	0.12	0.142	0.189
Female: immigrant	0.054	0.048	0.059	0.18	0.161	0.196
Immigrant (either parent)	0.086	0.079	0.091	0.272	0.244	0.272
Male: immigrant	0.070	0.062	0.071	0.214	0.186	0.208
Male: 15–24 years old	0.102	0.084	0.074	0.096	0.065	0.054
Male: 25–29 years old	0.317	0.286	0.242	0.299	0.28	0.231
Male: 30–34 years old	0.355	0.374	0.382	0.352	0.386	0.408
Male: 35–39 years old	0.177	0.189	0.226	0.195	0.196	0.232
Male: 40–44 years old	0.038	0.047	0.053	0.043	0.053	0.055
Male: 45 and older	0.011	0.021	0.024	0.016	0.022	0.022
Male: allophone	0.051	0.062	0.075	0.091	0.155	0.174
Male: francophone	0.867	0.87	0.854	0.03	0.055	0.05
Male: anglophone	0.083	0.065	0.067	0.88	0.788	0.774
Male: high school dropout	0.268	0.245	0.205	0.273	0.225	0.179
Male: high school diploma	0.163	0.155	0.158	0.145	0.163	0.157
Male: some post-secondary	0.439	0.463	0.476	0.43	0.456	0.487
Male: university degree	0.131	0.137	0.162	0.151	0.156	0.177
Live in urban area	0.765	0.762	0.762	0.826	0.82	0.835
Income: under C$19,999	0.262	0.261	0.331	0.191	0.178	0.242
Income: C$20,000–39,999	0.449	0.461	0.429	0.407	0.428	0.413
Income: C$40,000–59,999	0.227	0.209	0.182	0.308	0.289	0.248
Income: C$60,000–79,999	0.042	0.046	0.038	0.065	0.068	0.061
Income: C$80,000 and higher	0.019	0.022	0.019	0.03	0.036	0.033
Sum of weights	476435	468445	377825	610005	589105	510670

Notes: Each entry is the proportion of the weighted sample for each variable of each census file. For variable Had a child, we use periods from 1984 to 1985 for census 1986, from 1987 to 1988 for census 1991, and from 1994 to 1995 for census 1996. Observations are weighted and are rounded to the nearest multiple of 5 in Quebec or Ontario.[27] We also control for age, education level, mother tongue, and immigrant status of both the woman and her spouse. Table 3.2 shows proportions of the weighted sample for all variables of each census file.

3.5 Empirical Methods

To start, we replicate Milligan's (2005) difference-in-differences model to ensure continuity before examining heterogeneous responses and testing for permanent and transitory effects. After replicating Milligan's (2005) model with our data, we estimate the same model by sub-sampling different sibship sex compositions, income groups, maternal education, birth order, and immigrant status. We first estimate the following equation with Milligan's specification, and then proceed to run the same equation but with the abovementioned subgroups of the population:

$$\text{Had a Child}_{ijt} = \beta_0 + \beta_1 \text{Quebec}_j + \beta_2 \text{Census1996}_t + \beta_3 \text{Quebec}_j \times \text{Census1996}_t + X'_{ijt}\beta + \varepsilon_{ijt} \tag{3.1}$$

For equation (3.1), i indexes the individual females, j indexes jurisdictions, and t indexes time. The dependent variable indicates whether a child is born. Dummy variables are included to control for time effects, Census1996_t, and Quebec fixed effects, Quebec_j. The interaction of the two, $\text{Quebec}_j \times \text{Census1996}_t$, is our main variable of interest and accounts for any differential trend in fertility among residents of Quebec relative to those in Ontario. These models are estimated using probit regression and all standard errors (ε_{ijt}) are adjusted for heteroscedasticity.[28] Average marginal effects are reported to allow for easier interpretation of the estimates.[29] These marginal probabilities are interpreted as the marginal probability of having a child for a change in the independent variable of interest.[30]

[27] A household is located in an urban dwelling if it is located in a census metropolitan area (CMA), which is one or more municipalities with at least one hundred thousand people.

[28] In some instances we also utilized a triple-difference model, however we prefer the ease of interpretation provided by sub-sampling the difference-in-differences model. The triple-difference results match well with our preferred model. Results of the triple-difference are available upon request.

[29] Special care is taken into calculating average partial effects instead of partial effects evaluated at the mean. We observe individual's characteristics to calculate an individual probability and then average all those probabilities, as opposed to mean marginal effects, where the mean for each variable is plugged in to calculate a probability. We calculate the marginal probability using the method described in Ai and Norton (2003).

[30] The approach of using "probability of having a child" as the dependent variable is not new to this literature (see Cohen et al. 2013).

The variables included in X_{ijt} relate to the individual woman, her spouse, and her household. Age dummies signify whether the woman is between 15 to 24, 25 to 29, or 30 to 34 years of age, immigrant status, and, mother tongue.[31] Highest level of education is one of the subgroups we model for heterogeneous responses; we categorize education as high school dropout, high school diploma, some post-secondary, and a bachelor's degree or more. Similar variables are included for the spouse;[32] the only difference is age, for which the categories are 15 to 24, 25 to 29, 30 to 34, 35 to 39, 40 to 44, and 45 and older. Real annual family income excluding the woman's wage is categorized as under C$19,999, C$20,000 to C$39,999, C$40,000 to C$59,999, C$60,000 to C$79,999, and C$80,000 and over. We also account for the number of children already in the household: none, one, and two or more. A dummy variable is included to signify whether the household lives in an urban area.

To show how the ANC affected the timing of births we modify the outcome variable in equation (3.1) to be a binary indicator to signify two or more births within three years, two or more births within five years, and three or more births within five years; the probit estimates will show whether the ANC affected the timing of births and for which subgroups. It is important also to examine how the policy affected the total number of children born to each mother. For married or common-law women aged 35 to 39 from the 1991 and 2001 Censuses we estimate both linear and probit models similar to equation (1); the dependent variable in the linear model is the total number of children born to each woman and the outcome variable in the probit models is a binary indicator signifying that the woman had n children in total, where $n = 1$, 2, or 3 or more in each separate model.[33] In addition to examining the total number of children in a household, we further examine the sex composition of a three-child family. The dependent variable in each separate probit model is: had three sons, three daughters, one son and two daughters, and one daughter and two sons. Table 3.3 contains summary statistics from the 1991 and 2001 Censuses for married women aged 35 to 39 in both Ontario and Quebec.[34]

[31]The definition of immigrant in this case comes from the Census definition, which represents all individuals not born as a Canadian citizen.

[32]Nitsche et al. (2018) find evidence that it is important to also account for the male partner's education level as it also significantly predicts fertility.

[33]We also estimated $n = 4$ or more children and find that the results are similar to those for $n = 3$ or more.

[34]We limit our sample to women aged 35 to 39 because they are near the end of child-bearing, while still being young enough to have their children living at home. The Census only accounts for the number of children present in the household, thus if we include older women we may be missing

Table 3.3: Census summary statistics for married women 35–39 years old

	Quebec		Ontario	
	1991	2001	1991	2001
Number of children	1.753	1.746	1.875	1.819
Had one child	0.196	0.199	0.167	0.17
Had two children	0.445	0.427	0.441	0.44
Had three or more children	0.205	0.211	0.251	0.231
Female: allophone	0.075	0.106	0.199	0.249
Female: francophone	0.851	0.826	0.058	0.047
Female: anglophone	0.074	0.068	0.743	0.704
Female: high school dropout	0.259	0.167	0.212	0.161
Female: high school diploma	0.26	0.195	0.219	0.163
Female: some post-secondary	0.367	0.452	0.401	0.465
Female: university degree	0.114	0.186	0.168	0.211
Female: immigrant	0.083	0.097	0.257	0.282
Male: immigrant	0.1	0.1	0.29	0.271
Male: allophone	0.085	0.122	0.214	0.269
Male: francophone	0.844	0.812	0.055	0.045
Male: anglophone	0.071	0.066	0.731	0.686
Male: high school dropout	0.268	0.221	0.23	0.2
Male: high school diploma	0.168	0.169	0.136	0.14
Male: some post-secondary	0.402	0.43	0.429	0.459
Male: university degree	0.162	0.18	0.205	0.201
Live in urban area	0.754	0.767	0.827	0.861
Income: under C$19,999	0.228	0.258	0.162	0.209
Income: C$20,000–39,999	0.361	0.365	0.312	0.312
Income: C$40,000–59,999	0.272	0.236	0.322	0.269
Income: C$60,000–79,999	0.089	0.081	0.126	0.113
Income: C$80,000 and higher	0.05	0.06	0.078	0.097
Sum of weights	211,320	210,920	295,400	342,490

Notes: Each entry is the proportion of the weighted sample for each variable of each census file. Observations are weighted and are rounded to the nearest multiple of 5.

3.6 Results

3.6.1 Findings

Table 3.4 displays the average marginal effects for equation (1) with our entire sample as a replication exercise of Milligan (2005), but using the confidential data file. Our average marginal effects provide more accurate estimates and almost all are statistically significant at the one percent level. The first column displays controls with female characteristics, the second male characteristics, and the third family income and whether or not the family lives in an urban area. With all our variables included in

children that are no longer living at home.

the regression, the marginal effect of the interaction $\text{Quebec}_j \times \text{Census1996}_t$ displays a 1.8 percentage point increase in the probability of having a child. This translates to an 8.6 percent implied increase in the probability of having a child.[35] As a comparison Milligan (2005) estimates a 1.3 percentage point increase in the probability of having a child and the implied percentage increase is 8.7.

Table 3.4: Average marginal effects

	(a)	(b)	(c)
Census1996 × Quebec	0.0100	0.0201	0.0178
	(0.0017)	(0.0016)	(0.0015)
Census1996	-0.0042	-0.0133	0.0020
	(0.0011)	(0.0011)	(0.0010)
Quebec	0.0024	-0.0117	0.0003
	(0.0012)	(0.0015)	(0.0014)
One older child	0.4509	0.2314	0.1560
	(0.0016)	(0.0017)	(0.0014)
Two or more older children	0.1214	-0.0358	-0.0736
	(0.0017)	(0.0013)	(0.0012)
Female: 25–34 years old		0.2864	0.1174
		(0.0009)	(0.0010)
Female: immigrant		0.0755	0.0597
		(0.0014)	(0.0014)
Female: francophone		0.1276	0.1201
		(0.0015)	(0.0017)
Female: anglophone		0.1136	0.1107
		(0.0012)	(0.0013)
Female: high school diploma		0.0018	-0.0173
		(0.0012)	(0.0011)
Female: some post-secondary		-0.0177	-0.0353
		(0.0010)	(0.0010)
Female: university degree		-0.0790	-0.0888
		(0.0012)	(0.0012)
Male: 25–34 years old			0.0822
			(0.0021)
Male: 35–44 years old			0.0466
			(0.0022)
Male: 45 and older			-0.1171
			(0.0038)
Male: immigrant			-0.0511
			(0.0012)
Male: francophone			-0.1397
			(0.0012)
Male: anglophone			-0.1203
			(0.0011)
Male: high school diploma			-0.0081
			(0.0013)
Male: some post-secondary			-0.0050

[35]This calculation is based on the average marginal effect for the interaction term divided by the proportion of women that had a child in Quebec in our pre-policy period (1987-88), which was 0.207.

Table 3.4 (continued)

	(a)	(b)	(c)
			(0.0010)
Male: university degree			0.0006
			(0.0014)
Married			0.1947
			(0.0034)
Live in urban area			-0.0277
			(0.0009)
Income			0.0002
			(0.0000)
Pseudo-R-squared	0.0778	0.1789	0.2986
Number of observations	953630.0000	953630.0000	953630.0000

Notes: Dependent variable is Had a child. Robust standard errors are in parenthesis. Observations are rounded to the nearest multiple of 5

Table 3.5 shows the average marginal effect of the ANC policy on having a child from equation (1) with each panel displaying a different group of interest (birth parity, sibship sex composition, income group, education level, and immigrant status.). Each column represents the sub-sample for which a separate probit regression is estimated. The rows in each panel display the average marginal effect for the interaction term $\text{Quebec}_j \times \text{Census1996}_t$, its standard error, the implied percentage increase in the probability of having a child, the probability of having a child based on a representative woman, the pseudo R-squared from the probit model, the pre-policy and during policy rate, and the number of observations used in the regression. The implied percentage increase in probability of having a child is calculated by dividing the average marginal effect of the ANC policy (the interaction term for $\text{Quebec}_j \times \text{Census1996}_t$) by the proportion of women in each sub-sample that had a child in Quebec in our pre-policy period (1987-1988). The probability of having a child based on a representative woman is calculated using the probit coefficients, and the representative woman is described in the notes section of Table 3.5. The pre-policy and during policy rate are the proportion of our subsample that had a child in Quebec in the respective time period.

In Panel A of Table 3.5 , we find a large and statistically significant effect on birth order, specifically for families that already have two children: the estimates imply a twenty-three percent increase in the probability of having their third or higher child. The baby bonus also increases the implied marginal probability for first and second children by ten and three percent, respectively; however it is clear the baby bonus supported higher birth order children most by providing a very generous baby bonus ($8000 for third or higher order children). Women with a previous child have a very high probability of having a second (42.2 percent based on our representative woman), so the baby bonus was more likely to encourage third or higher birth order children. Laroque and Salanié (2008) also find evidence that first and third births are responsive to financial incentives in their examination of France's Allocation Parental d'Education (APE). Most people that already have their first child are also going to have a second, whereas cash incentives either encourage first-time parents or parents of two to try for a third. Panel B indicates that a gender preference for sons is present. We find that there is a statistically significant effect of the ANC policy for families that already have a son, but not for those who already have a daughter. The implied percentage increase is much higher for those with a previous son, demonstrating that families with a strong preference for a son and a previous daughter were planning to have another child regardless of the baby bonus, whereas families with a son were more encouraged to have another child. These results are in line with studies that find a stronger preference for sons (see Almond et al. (2013), Dahl and Moretti (2008)).

Table 3.5: Average marginal effects

Panel A: subsamples by birth order			
	No older children	One older child	Two or more older children
Average marginal effect (Quebec×Census1996)	0.0179	0.0127	0.0288
	(0.0045)	(0.0068)	(0.0050)
Implied percentage increase in probability of having a child	0.1040	0.0320	0.2330
Probability of having a child[a]	0.3640	0.4220	0.1250
Pseudo-R-squared	0.0536	0.0215	0.0310
Pre-policy rate	0.1720	0.3910	0.1230
During policy rate	0.2600	0.4120	0.1500
Number of observations	213010	83940	93080
Panel B: subsamples by gender of first child			
	Son	Daughter	
Average marginal effect (Quebec×Census1996)	0.0213	0.0023	
	(0.0094)	(0.0096)	
Implied percentage increase in probability of having a child	0.0560	0.0060	
Probability of having a child[a]	0.4250	0.4180	
Pseudo-R-squared	0.0218	0.0237	
Pre-policy rate	0.3830	0.3860	
During policy rate	0.4110	0.4040	
Number of observations	43770	41770	
Panel C: subsamples by gender of previous two children			
	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0381	0.0485	0.0357

Table 3.5 (continued)

	(0.0106)	(0.0165)	(0.0161)	
Implied percentage increase in probability of having a child	0.3380	0.3300	0.2330	
Probability of having a child[a]	0.1180	0.1530	0.1320	
Pseudo-R-squared	0.0328	0.0281	0.0306	
Pre-policy rate	0.1130	0.1470	0.1530	
During policy rate	0.1430	0.1750	0.1740	
Number of observations	35770	17975	16255	
Panel D: subsamples by income group				
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999
Average marginal effect (Quebec×Census1996)	0.0057	0.0193	0.0312	0.0233
	(0.0052)	(0.0042)	(0.0060)	(0.0129)
Implied percentage increase in probability of having a child	0.0160	0.0950	0.1360	0.0910
Probability of having a child[a]	0.2040	0.3020	0.3420	0.3410
Pseudo-R-squared	0.0722	0.0654	0.0710	0.0823
Pre-policy rate	0.1770	0.2030	0.2300	0.2550
During policy rate	0.2470	0.2690	0.2900	0.2950
Number of observations	97410	168170	92300	21300
Panel E: subsamples by women's education				
	High school dropout	High school diploma	Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	0.0117	0.0121	0.0231	0.0263
	(0.0065)	(0.0065)	(0.0039)	(0.0068)
Implied percentage increase in probability of having a child	0.0530	0.0520	0.1170	0.1520
Probability of having a child[a]	0.2540	0.2860	0.3090	0.3380
Pseudo-R-squared	0.0629	0.0693	0.0740	0.1211
Pre-policy rate	0.2210	0.2340	0.1970	0.1730
During policy rate	0.2570	0.2500	0.2740	0.2750
Number of observations	72545	72775	185170	59540
Panel F: subsamples by immigration status				
	Non-immigrant (both parents)		Immigrant (either or both parents)	
Average marginal effect (Quebec×Census1996)	0.0171		0.0233	
	(0.0030)		(0.0083)	
Implied percentage increase in probability of having a child	0.0840		0.0970	
Probability of having a child[a]	0.2870		0.3120	
Pseudo-R-squared	0.0686		0.0560	
Pre-policy rate	0.2030		0.2410	
During policy rate	0.2640		0.3030	
Number of observations	321245		68780	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5. The pre-policy and during policy rate are the proportion of our subsample that had a child in the respective time period

Table 3.5 (continued)

[a] The probability of having a child is calculated based on a representative woman who is married, francophone, lives in an urban area in Quebec, 30–34 years old, during the policy period; (in panel A) is a non-immigrant with some post-secondary education and has no previous children; (in panel B) is a non-immigrant with a family income between $20,000–$40,000 and has no previous children; (in panel C) is a non-immigrant with some post-secondary education, with a family income between $20,000–$40,000; (in panel D) is a non-immigrant with some post-secondary education, with a family income between $20,000–$40,000, and already has two previous children; and (in panel E) has some post-secondary education, with a family income between $20,000–$40,000, and has no previous children

Panel C further delves into sex preference with the third child by controlling for the sex of the previous two children. We find that the baby bonus provided the same incentive to have a third child for parents with two previous sons or with a son and a daughter (by 33.0 percent and 33.8 percent respectively) but somewhat less for parents with two daughters (23.3 percent). That suggests that the baby bonus encouraged more births from parents who otherwise would have stopped at two: prior to the policy, parents with two daughters were more inclined to have a third child than parents with both a son and daughter. In Quebec during our pre-policy period (1987-1988) the percent of parents with two previous daughters that had a third child was 15.3 percent, whereas the percent of parents with a previous son and daughter was only 11.3 percent. This follows well-documented empirical evidence that parents are more likely to go for a third child when they have two previous daughters Angrist et al., 2010.

In panel D we observe a hump shape response to the ANC by income groups. Interestingly, the lowest (under C$20,000) and highest (over C$80,000) income groups' response is not statistically significant, and the coefficients are very small. The second lowest (C$20,000-C$40,000) and highest (C$60,000-C$80,000) income groups both have an implied 9.5 percent increase in the probability of having a child that is statistically significant. Finally, the mid-income group (C$40,000-C$60,000) has the largest response with an implied increase of almost 14 percent in the probability of having a child; this result is statistically significant at the one percent level. Once the policy is implemented, we see Quebecois women with a higher probability of having a child, and the rise is predominantly in the mid-income range.[36]

Panel E shows the response by level of education of the women. All the results are statistically significant; however we observe that women with a high school diploma or less have a five percent increase in the implied probability of having a child due

[36]The representative female used to calculate the probability of having a child is a married non-immigrant francophone woman who is 30-34 years old, with some post-secondary education, lives in an urban area, and has no previous children. These characteristics are chosen as they are the most common female we encounter in our sample and thus make the most general comparison.

to the baby bonus and an even greater response among women with more education: the implied percentage increase is twice as great for women with some post-secondary education and three times as great for those with a bachelor's degree or higher. This is consistent with recent work that suggests highly educated women are opting for more children (see Shang and Weinberg, 2013). Moreover, the probability of having a child follows the same positive gradient across female education levels.[37]

Finally, in Panel F we examine the response based on immigrant status. The results for both immigrant and non-immigrants are positive statistically significant at the one percent level and suggest that the immigrant response is only slightly greater. When both parents are non-immigrants there is an implied eight percent increase in the probability of having a child; when either or both are immigrants it is almost ten percent. Both have around a thirty percent likelihood of having a child.

Table 3.6: Average marginal effects of ANC on child spacing by subsample

Panel A: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Dependent variable: had 2 or more kids in 3 years					
Average marginal effect (Quebec×Census1996)	0.00730	0.01370	0.00760	0.00530	0.01550
	(0.00460)	(0.00460)	(0.00350)	(0.00510)	(0.00660)
Implied percentage increase in probability	0.20500	0.39100	0.59200	0.40000	1.08600
Dependent variable: had 2 or more kids in 5 years					
Average marginal effect (Quebec×Census1996)	0.03370	0.04730	0.00950	0.02140	0.02110
	(0.00840)	(0.00830)	(0.00740)	(0.01150)	(0.01200)
Implied percentage increase in probability	1.43300	1.82600	0.87300	1.62900	2.10700
Dependent variable: had 3 or more kids in 5 years					
Average marginal effect (Quebec×Census1996)	0.00910	0.01170	0.00060	-0.0004	0.01100
	(0.00330)	(0.00360)	(0.00240)	(0.00340)	(0.00690)
Implied percentage increase in probability	0.38700	0.45200	0.05500	-3.0%	1.09800
Number of observations	31030	32535	18470	9470	8495
Panel B: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher

[37]Since younger women are likely to return to school, as a robustness check we estimate our specification considering only women over 25 years of age and results do not change significantly.

Table 3.6 (continued)

Dependent variable: had 2 or more kids in 3 years					
Average marginal effect (Quebec×Census1996)	0.00500	0.00810	0.01350	0.01520	0.01940
	(0.00210)	(0.00210)	(0.00330)	(0.00730)	(0.01040)
Implied percentage increase in probability	0.17700	0.29000	0.37600	0.37200	0.45100
Dependent variable: had 2 or more kids in 5 years					
Average marginal effect (Quebec×Census1996)	0.01610	0.02380	0.03380	0.04170	0.03850
	(0.00340)	(0.00340)	(0.00520)	(0.01140)	(0.01710)
Implied percentage increase in probability	0.19100	0.24500	0.28200	0.28600	0.24700
Dependent variable: had 3 or more kids in 5 years					
Average marginal effect (Quebec×Census1996)	0.00410	0.00440	0.00500	0.01270	0.00520
	(0.00130)	(0.00120)	(0.00190)	(0.00520)	(0.00560)
Implied percentage increase in probability	0.59300	0.73900	0.83300	1.41100	0.37300
Number of observations	137785	168170	92300	21295	10845
Panel C: subsamples by women's education					
	High school dropout	High school diploma	Some post-secondary	Bachelor degree or higher	
Dependent variable: had 2 or more kids in 3 years					
Average marginal effect (Quebec×Census1996)	0.00890	0.00780	0.00890	0.00520	
	(0.00350)	(0.00330)	(0.00200)	(0.00350)	
Implied percentage increase in probability	0.27100	0.22200	0.30700	0.18600	
Dependent variable: had 2 or more kids in 5 years					
Average marginal effect (Quebec×Census1996)	0.00400	0.02660	0.02690	0.02290	
	(0.00550)	(0.00550)	(0.00330)	(0.00540)	
Implied percentage increase in probability	0.03400	0.22700	0.28700	0.28600	
Dependent variable: had 3 or more kids in 5 years					
Average marginal effect (Quebec×Census1996)	0.00340	0.00410	0.00530	0.00230	
	(0.00210)	(0.00190)	(0.00120)	(0.00170)	
Implied percentage increase in probability	0.41900	0.51500	0.88500	0.45800	
Number of observations	72540	72780	185165	59540	
Panel D: subsamples by immigrant status					
		Non-immigrant (both parents)	Immigrant (either or both parents)		
Dependent variable: had 2 or more kids in 3 years					

Table 3.6 (continued)

Average marginal effect (Quebec×Census1996)	0.00770	0.01060
	(0.00150)	(0.00420)
Implied percentage increase in probability	0.25800	0.28600
Dependent variable: had 2 or more kids in 5 years		
Average marginal effect (Quebec×Census1996)	0.02200	0.02520
	(0.00250)	(0.00680)
Implied percentage increase in probability	0.21800	0.21200
Dependent variable: had 3 or more kids in 5 years		
Average marginal effect (Quebec×Census1996)	0.00440	0.00400
	(0.00090)	(0.00230)
Implied percentage increase in probability	0.73600	0.44900
Number of observations	321245.00000	68780.00000

Notes: The implied percentage increase is calculated by dividing the average marginal effect from the Quebec pre-policy dependent variable by each respective subsample. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5.

Table 3.6 examines the impact of sibship sex composition, family income, mother's education, and immigrant status on the spacing of children to assess whether the baby bonus encouraged families to have their children closer together. In the first panel, in general, the baby bonus encouraged a rise in tempo. Specifically, we see that families with a daughter are even more inclined to space children closer together, and this result is statistically significant across all three separate regressions. When examining the gender of two previous children, it is the family that already has two daughters that is spacing their children closer together. In Panel B we find as family income increases more children are spaced closer together; the results are statistically significant. In Panel C, all the marginal effects by mother's education are positive and mostly statistically significant; that suggests that the baby bonus encouraged parents to space their children closer together. Panel D continues to show that the baby bonus affected non-immigrant and immigrant families similarly. The final row of Panel D does show a much larger implied percentage increase in the probability of non-immigrants having three or more children in five years, but it is based on relatively few observations.

Table 3.7: Average marginal effects of ANC on completed fertility

Panel A: linear model	
Dependent variable: total number of children	
Average marginal effect (Quebec×Census2001)	0.0417
	(0.0099)
Implied percentage increase	2.40%
Number of observations	208,560
Panel B: probit model	
Dependent variable: family had 1 child	
Average marginal effect (Quebec×Census2001)	0.0008
	(0.0034)
Implied percentage increase in probability	0.40%
Dependent variable: family had 2 children	
Average marginal effect (Quebec×Census2001)	-0.0176
	(0.0044)
Implied percentage increase in probability	-4.0%
Dependent variable: family had 3 or more children	
Average marginal effect (Quebec×Census2001)	0.0208
	(0.0038)
Implied percentage increase in probability	10.20%
Number of observations	208,560
Panel C: 3-child family formation (probit model)	
Dependent variable: had 3 sons	
Average marginal effect (Quebec×Census2001)	0.0023
	(0.0016)
Implied percentage increase in probability	9.10%
Dependent variable: had 3 daughters	
Average marginal effect (Quebec×Census2001)	0.0015
	(0.0014)
Implied percentage increase in probability	7.00%
Dependent variable: had 1 son and 2 daughters	
Average marginal effect (Quebec×Census2001)	0.002
	(0.0023)
Implied percentage increase in probability	3.40%
Dependent variable: had 1 daughter and 2 sons	
Average marginal effect (Quebec×Census2001)	0.0093
	(0.0024)
Implied percentage increase in probability	15.70%
Number of observations	195,620

Notes: The implied percentage increase is calculated by dividing the average marginal effect from the Quebec pre-policy dependent variable. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5.

Using the 1991 and 2001 Censuses, Table 3.7 asks whether the ANC increased fertility and thus had a permanent effect. Panel A displays the results from a linear model where the dependent variable is the total number of children each family has. Here we see that there was a positive statistically significant effect which implied a 2.4 percent increase in the total number of children. To further examine how the ANC affected the total number of children born per family, probit models are estimated. The results in Panel B suggest that the baby bonus had a statistically insignificant and economically negligible effect on the probability of having one child and a negative effect on having two children.[38] However, the ANC policy had a positive, large, and statistically significant effect on families with three or more children. However, the ANC policy had a positive, large, and statistically significant effect on families with three or more children. Specifically, there was a 10.2 percent increase in the probability of having a family with three or more kids due to the Quebec baby bonus. Since we observe that the ANC policy had a large impact on three-child families, we examine which of these family formations had the greatest increase due to the policy. Panel C shows that there was a statistically significant sixteen percent increase in the number of three-child families that had one daughter and two sons due to the ANC policy.

Table 3.8 follows the same probit model as Panel B of Table 3.7 but subgroups by family income, mother's education, and immigrant status. Here again we see that the ANC was mostly statistically insignificant for one child families, has a negative coefficient on the two child household, and had a major impact on increasing family size to three or more children.

Table 3.8: Average marginal effects of ANC on number of children by subsample

Panel A: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Dependent variable: family had 1 child					
Average marginal effect (Quebec×Census2001)	0.0082	0.0006	-0.0062	0.01	-0.0197
	(0.0074)	(0.0061)	(0.0066)	(0.0107)	(0.0129)
Implied percentage increase in probability	3.90%	0.30%	-3.3%	5.70%	-12.2%
Dependent variable: family had 2 children					
Average marginal effect (Quebec×Census2001)	-0.0137	-0.0189	-0.0128	-0.0370	-0.0148
	(0.0091)	(0.0076)	(0.0086)	(0.0144)	(0.0177)
Implied percentage increase in probability	-3.4%	-4.2%	-2.7%	-7.8%	-3.3%
Dependent variable: family had 3 or more children					
Average marginal effect (Quebec×Census2001)	0.0037	0.0216	0.0161	0.0375	0.0411
	-0.0081	-0.0064	-0.0072	-0.0123	-0.0158
Implied percentage increase in probability	1.70%	11.00%	8.20%	17.70%	16.30%
Number of observations	43,920	69,020	57,830	21,900	15,350

[38] Households' response to having two children is negative since they are likely moving to a family with three children given the large cash incentive.

Table 3.8 (continued)

Panel B: subsamples by women's education				
	High school dropout	High school diploma	Some post-secondary	Bachelor degree or higher
Dependent variable: family had 1 child				
Average marginal effect (Quebec×Census2001)	0.0152	-0.0010	-0.0061	-0.0193
	(0.0084)	(0.0078)	(0.0055)	(0.0090)
Implied percentage increase in probability	7.80%	-0.5%	-3.1%	-9.6%
Dependent variable: family had 2 children				
Average marginal effect (Quebec×Census2001)	-0.0502	-0.0239	-0.0111	0.0143
	(0.0108)	(0.0101)	(0.0070)	(0.0113)
Implied percentage increase in probability	-11.2%	-5.0%	-2.4%	3.60%
Dependent variable: family had 3 or more children				
Average marginal effect (Quebec×Census2001)	0.0387	0.0362	0.0163	0.0373
	(0.0098)	(0.0086)	(0.0060)	(0.0091)
Implied percentage increase in probability	16.40%	19.20%	8.20%	20.10%
Number of observations	35,505	40,130	84,860	35,130
Panel C: subsamples by immigrant status				
		Non-immigrant (both parents)		Immigrant (either or both parents)
Dependent variable: family had 1 child				
Average marginal effect (Quebec×Census2001)		0.0033		-0.0090
		(0.0041)		(0.0088)
Implied percentage increase in probability		1.60%		-5.1%
Dependent variable: family had 2 children				
Average marginal effect (Quebec×Census2001)		-0.0202		-0.0201
		(0.0052)		(0.0115)
Implied percentage increase in probability		-4.4%		-4.6%
Dependent variable: family had 3 or more children				
Average marginal effect (Quebec×Census2001)		0.0211		0.0334
		(0.0043)		(0.0103)
Implied percentage increase in probability		10.80%		12.30%
Number of observations		146,450		49,170

Notes: The implied percentage increase is calculated by dividing the average marginal effect from the Quebec pre-policy dependent variable by each respective subsample. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5.

3.6.2 Sensitivity Analysis

As a robustness check we re-estimate the same specification for Tables 3.5, 3.6 and 3.7 using a linear probability model instead of a probit model; we find similar results. As a second robustness check we re-estimate Table 3.5 without controlling for male characteristics since, as a consequence of assortative mating, they may be highly correlated with income and the female's characteristics. The exclusion of spousal characteristics does not alter our findings. These estimates can be found in the Appendix section under Table 3.9. Since we previously excluded families that had an income of zero from our sample, we now include these families back in. In Table 3.10 we see that our hump shape result for family income still holds. Low income families respond far less to the policy than mid-income families. Specifically, the coefficient estimate for family income between $0 and $20,000 decreases from 0.0057 to 0.0028 illustrating that the poorest of the poor are responding even less.

We also examine single females, previously excluded from our sample, in Table 3.11. We are able to divide the population of single females by marital status, which are those that have never been married versus those that are separated from a previous marriage. The baby bonus has a statistically insignificant positive effect on separated women, and a statistically significant negative effect on females that have never been married. Table 3.11 also shows the pre-policy rate of having a child in Quebec and these values are very small; the baby bonus was utilized by married women, and not encouraging new single mothers.

Next, we use 1984 and 1985 from the 1986 census file as our control group, instead of 1987 and 1988 from the 1991 census file. The results, found in Table A5, are substantially unchanged qualitatively, but exact estimates do vary because of the substantial decline in fertility in the mid-1980s. For example, we find the same hump shape by income but it is shifted up because the new control group (1984-1985) had fewer children, thereby creating a larger difference from the treatment group (1994-1995).

Furthermore, to minimize the cultural dissimilarity between our treated and control group, we conduct the following exercise: we estimate our model using only households living near the border of the two provinces. We find Table A6 results are qualitatively similar.[39]

[39] The sample size drops to 90,000 households. Also, Quebec has almost 4 times the number of observations than Ontario. Thus, this is not our preferred specification. The CMAs we selected

We re-estimate Table 3.5 with a three-year and five-year window inside and outside the policy to ensure our two-year window from all previous regressions is reliable. In the three-year window (Table 3.15) we use 1986 to 1988 as the pre-policy window and 1993 to 1995 as the within-policy window. For the five-year window (Table 3.14) we examine 1984 to 1988 versus 1991 to 1995. Qualitatively the results are the same and statistically significant with the same hump shape response in income, as well as the same heterogeneous responses in parity, sibship sex composition, and education.

Since immigrants may respond differently to the baby bonus, we exclude them from the sample and re-estimate the model. The results are shown in Table 3.16. In Table 3.17, we limit the sample to only females aged 25-34 since females younger than 25 are likely still in school. We find the response is weaker at the margin, but is consistent with our birth CDF findings; more females gave birth at younger ages under the ANC policy. These robustness checks confirm that the baby bonus did create heterogeneous responses among women. As a final check, we also use the exact match method that stratifies females with the same characteristics, and then perform a difference-in-differences calculation across time (pre-policy and within-policy) and across groups (Quebec and Ontario). This method relaxes the assumptions on global common trends and model dependence. We match females by birth parity, income group, education level, and age group. Each unique grouping forms a stratum. In this case we create 180 strata (3 parity groups $\times$ 5 income groups $\times$ 4 education levels $\times$ 3 age groups).[40] For each stratum we calculate the difference in having a child between Quebec and Ontario females as well as pre-policy and within-policy periods. This difference-in-differences calculation results in a hump shape response across income groups, confirming our earlier findings.[41]

3.7 Conclusion

When we examine the impact of the ANC on fertility by birth order we find a strong increase in the probability of having a third child or higher order. We are aware that these results are due to the specific payment structure of the ANC. From May 1992

are: Temiskming Shores, North Bay, Petawawa, Pembroke, Hawkesbury, Cornwall, Rouyn-Noranda, Lachute, Salaberry-de-Valleyfield, Val-d'Or, and Amos.

[40] We drop 12 strata because they contain less than 5 observations for each province and each period.

[41] Results available upon request.

until the cancellation of the policy in September 1997, the transfer payments were C$500 for the first child, C$1,000 for the second, and C$8,000 for the third child or higher. Had the payment structure provided a constant amount regardless of parity, the estimates for third or higher parity children would not be as large. The Quebec government continuously increased the transfer payment for third or higher children, from C$3,000 to C$8,000, demonstrating that they were also aware that families with two children already present in the household require a larger income transfer to induce them to have a third child.

North American parents prefer to have one-of-each gender, with a secondary preference for sons (Williamson, 1983). Interestingly, parents with two previous sons, or a previous son and daughter, were more inclined to have a third child after the ANC was implemented. This illustrates that parents who were more likely to stop at two children were successfully encouraged by the ANC to have another child. These results provide strong evidence to suggest that Quebec's baby bonus did in fact accomplish its goal of increasing fertility, while simultaneously alleviating the gender preferences of parents.

The heterogeneous responses we find suggest that baby bonuses do work. Pro-natalist policies can encourage household births by targeting the subgroups whose fertility decisions are highly responsive to cash incentives. For example, when examining the heterogeneous response of the ANC by income group, we find a hump shape result that is robust to many different specifications. Interestingly, Becker and Tomes (1976) model a U-shaped path for the desired number of children as income rises. This model predicts that a negative exogenous shock in the price of children would have low-income mothers spend extra income on children they already have rather than having more children. Moreover, the amount of the transfer may not be enough for high income individuals to be induced to have another child. Mid-income families seem poised to take advantage of a baby bonus, and if structured strategically pro-natalist policies can increase higher parity births. Furthermore, we find that highly educated women are more likely to participate in a baby bonus than less educated women. This encourages the reduction of the fertility rate disparity that is related to maternal education.

Moreover, we are able to observe the completed fertility rates of many cohorts that were exposed to the ANC. We see that, in addition to a transitory effect where women were having their children closer together, there was also an increase in completed fertility of women aged 15 to 39, illustrating that the pro-natalist policy does have a permanent effect on fertility in Quebec. We find that among 3-child households the baby bonus was able to create more one daughter-two son families then other sibship sex compositions. Pro-natalist policies, if structured correctly, could cost-effectively increase fertility and alleviate the immense concern of below-replacement rates for developed nations. Furthermore, pro-natalist policies can also diminish gender preferences by incentivizing parents to have more children.

3.8 Bibliography

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3.A Appendix

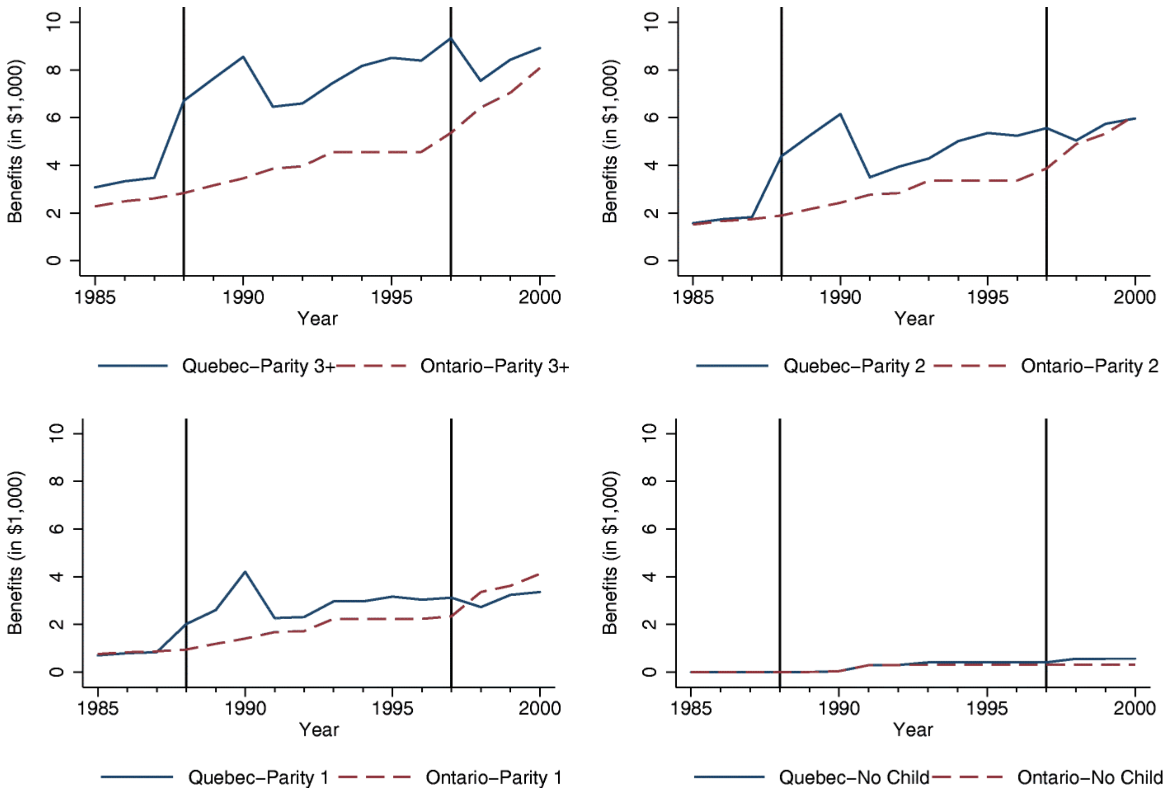

Figure 3.4: Total family benefit for household income of $20,000. The first vertical bar signifies the start of the ANC policy in May 1988, and the second vertical bar signifies the end of the policy in September 1997. Source: Milligan (2016a), Canadian Tax and Credit Simulator. Database, software, and documentation, version 2016-2

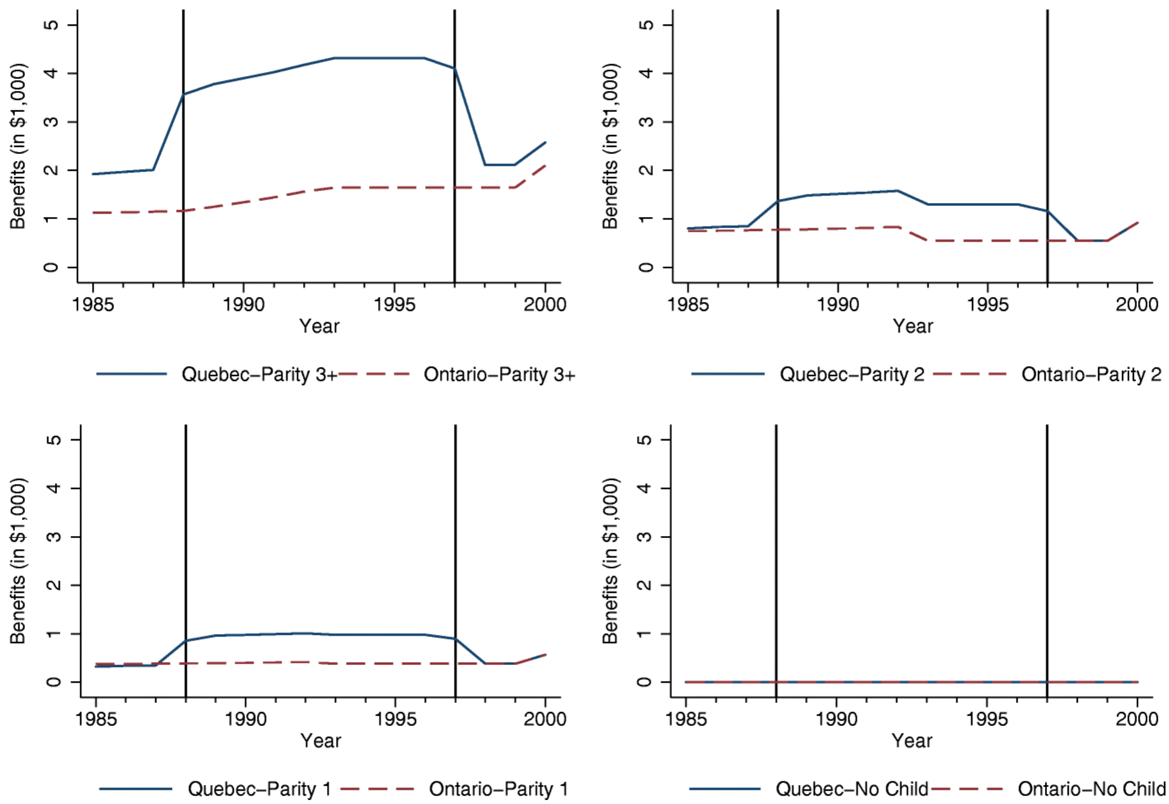

Figure 3.5: Total family benefit for household income of $60,000. The first vertical bar signifies the start of the ANC policy in May 1988, and the second vertical bar signifies the end of the policy in September 1997. Source: Milligan (2016a), Canadian Tax and Credit Simulator. Database, software, and documentation, version 2016-2

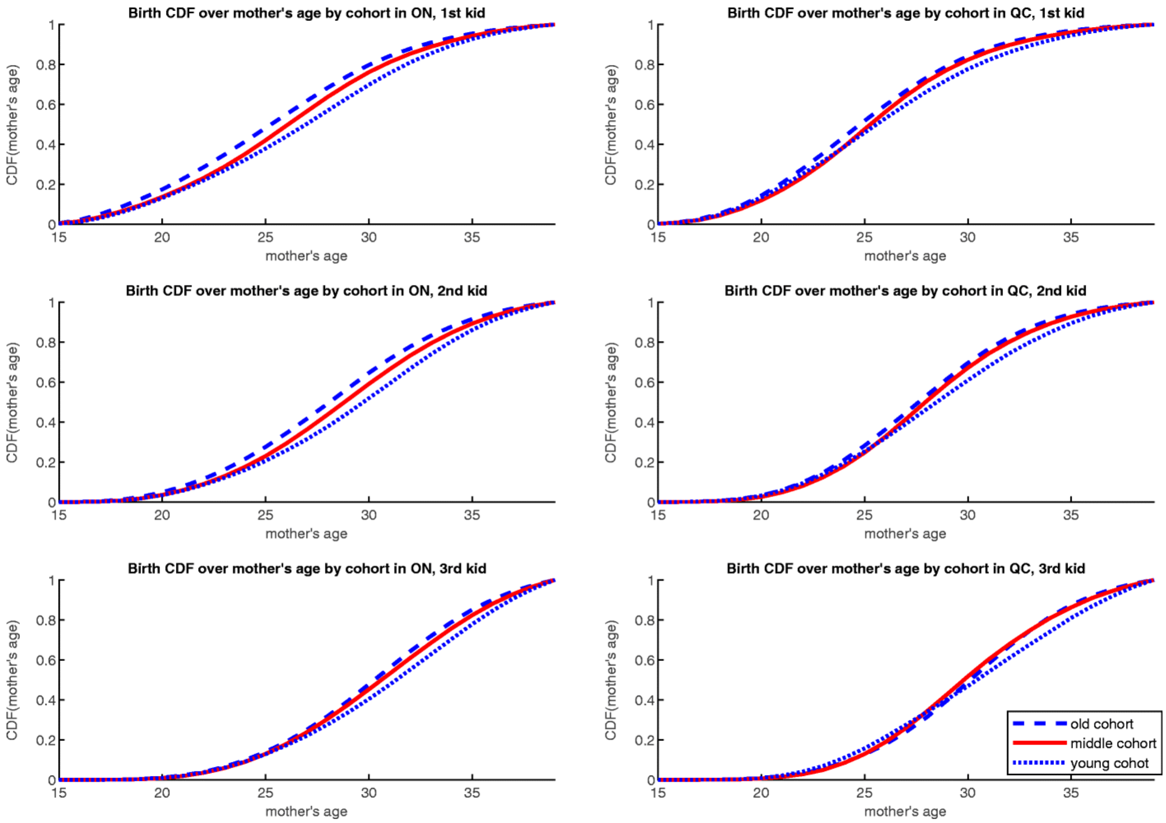

Figure 3.6: Birth cumulative distribution function by mother's age, cohorts aged 15–39. Birth Vital Statistics source. The "old cohort" was born between 1959 and 1962 and aged 26–38 during the policy; the "middle cohort" was born between 1963 and 1968 and aged 20–34 during the policy; and the "young cohort" was born between 1969 and 1972 and aged 16–28 during the policy

Table 3.9: Average marginal effects of ANC on child birth—excluding male characteristics

Panel A: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Dependent variable: family had 1 child					
Average marginal effect (Quebec×Census2001)	0.0082	0.0006	-0.0062	0.01	-0.0197
	(0.0074)	(0.0061)	(0.0066)	(0.0107)	(0.0129)
Implied percentage increase in probability	3.90%	0.30%	-3.3%	5.70%	-12.2%
Dependent variable: family had 2 children					
Average marginal effect (Quebec×Census2001)	-0.0137	-0.0189	-0.0128	-0.0370	-0.0148
	(0.0091)	(0.0076)	(0.0086)	(0.0144)	(0.0177)
Implied percentage increase in probability	-3.4%	-4.2%	-2.7%	-7.8%	-3.3%
Dependent variable: family had 3 or more children					
Average marginal effect (Quebec×Census2001)	0.0037	0.0216	0.0161	0.0375	0.0411
	-0.0081	-0.0064	-0.0072	-0.0123	-0.0158
Implied percentage increase in probability	1.70%	11.00%	8.20%	17.70%	16.30%
Number of observations	43,920	69,020	57,830	21,900	15,350
Panel B: subsamples by women's education					
	High school dropout	High school diploma	Some post-secondary	Bachelor degree or higher	
Dependent variable: family had 1 child					
Average marginal effect (Quebec×Census2001)	0.0152	-0.0010	-0.0061	-0.0193	
	(0.0084)	(0.0078)	(0.0055)	(0.0090)	
Implied percentage increase in probability	7.80%	-0.5%	-3.1%	-9.6%	
Dependent variable: family had 2 children					
Average marginal effect (Quebec×Census2001)	-0.0502	-0.0239	-0.0111	0.0143	
	(0.0108)	(0.0101)	(0.0070)	(0.0113)	
Implied percentage increase in probability	-11.2%	-5.0%	-2.4%	3.60%	
Dependent variable: family had 3 or more children					
Average marginal effect (Quebec×Census2001)	0.0387	0.0362	0.0163	0.0373	
	(0.0098)	(0.0086)	(0.0060)	(0.0091)	
Implied percentage increase in probability	16.40%	19.20%	8.20%	20.10%	
Number of observations	35,505	40,130	84,860	35,130	
Panel C: subsamples by immigrant status					

Table 3.9 (continued)

	Non-immigrant (both parents)	Immigrant (either or both parents)
Dependent variable: family had 1 child		
Average marginal effect (Quebec×Census2001)	0.0033	-0.0090
	(0.0041)	(0.0088)
Implied percentage increase in probability	1.60%	-5.1%
Dependent variable: family had 2 children		
Average marginal effect (Quebec×Census2001)	-0.0202	-0.0201
	(0.0052)	(0.0115)
Implied percentage increase in probability	-4.4%	-4.6%
Dependent variable: family had 3 or more children		
Average marginal effect (Quebec×Census2001)	0.0211	0.0334
	(0.0043)	(0.0103)
Implied percentage increase in probability	10.80%	12.30%
Number of observations	146,450	49,170

Notes: Dependent variable is Had a child. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5.

Table 3.10: Average marginal effects of ANC on child birth—including zero income households

Panel A: subsamples by birth order					
	No older children		One older child	Two or more older children	
Average marginal effect (Quebec×Census1996)	0.0225		0.0131	0.0294	
	(0.0035)		(0.0066)	(0.0048)	
Number of observations	240,140		89,385	100,875	
Panel B: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0218	0.0041	0.0268	0.0359	0.0345
	(0.0092)	(0.0094)	(0.0074)	(0.0116)	(0.0122)
Number of observations	45,755	43,630	38,360	19,240	17,425
Panel C: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Average marginal effect (Quebec×Census1996)	0.0028	0.0193	0.0312	0.0233	-0.0006
	(0.0042)	(0.0042)	(0.0060)	(0.0129)	(0.0190)
Number of observations	137,785	168,170	92,300	21,295	10,845
Panel D: subsamples by women's education					
	High school dropout	High school diploma		Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	0.0112	0.0139		0.0252	0.0295
	(0.0058)	(0.0061)		(0.0037)	(0.0066)
Number of observations	87,905	80,015		199,850	62,620
Panel E: subsamples by immigration status					
		Non-immigrant (both parents)		Immigrant (either or both parents)	
Average marginal effect (Quebec×Census1996)		0.0189		0.0304	
		(0.0028)		(0.0077)	
Number of observations		351,750		78,645	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.11: Average marginal effects of ANC on child birth—single females (84 to 88 vs. 91 to 95)

Panel A: subsamples by marital status				
	Never married	Separated		
Average marginal effect (Quebec×Census1996)	-0.0019	0.0019		
	(0.0008)	(0.0071)		
Quebec pre-policy rate	1.80%	13.30%		
Number of observations	482,330	38,495		
Panel B: subsamples by birth order				
	No older children	One older child	Two or more older children	
Average marginal effect (Quebec×Census1996)	-0.0011	0.0221	0.0392	
	(0.0007)	(0.0102)	(0.0107)	
Quebec pre-policy rate	1.80%	15.70%	7.80%	
Number of observations	478,820	24,405	17,600	
Panel C: subsamples by women's education				
	High school dropout	High school diploma	Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	0.0037	-0.0002	0.0007	-0.0012
	(0.0017)	(0.0023)	(0.0014)	(0.0017)
Quebec pre-policy rate	3.60%	2.50%	2.10%	1.00%
Number of observations	181,070	80,440	201,825	57,490
Panel D: subsamples by immigration status				
	Non-immigrant (both parents)	Immigrant (either or both parents)		
Average marginal effect (Quebec×Census1996)	0.0018	0.0010		
	(0.0010)	(0.0034)		
Quebec pre-policy rate	2.70%	1.60%		
Number of observations	398,090	122,735		

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.12: Average marginal effects of ANC on child birth—84 to 85 vs. 94 to 95

Panel A: subsamples by birth order					
	No older children		One older child		Two or more older children
Average marginal effect (Quebec×Census1996)	0.0286		0.0284		0.0195
	(0.0041)		(0.0065)		(0.0043)
Number of observations	187,900		93,420		118,480
Panel B: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0428	0.0127	0.0310	0.0252	0.0163
	(0.0091)	(0.0093)	(0.0091)	(0.0146)	(0.0143)
Number of observations	47,830	45,345	44,845	22,525	20,485
Panel C: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Average marginal effect (Quebec×Census1996)	0.0128	0.0225	0.0386	0.0518	0.0020
	(0.0053)	(0.0042)	(0.0058)	(0.0129)	(0.0194)
Number of observations	100,650	168,310	98,630	21,080	10,080
Panel D: subsamples by women's education					
	High school dropout	High school diploma		Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	0.0079	0.0345		0.0189	0.0420
	(0.0062)	(0.0064)		(0.0041)	(0.0077)
Number of observations	89,370	78,480		176,670	55,290
Panel E: subsamples by immigration status					
		Non-immigrant (both parents)		Immigrant (either or both parents)	
Average marginal effect (Quebec×Census1996)		0.0259		0.0179	
		(0.0030)		(0.0081)	
Number of observations		324,220		74,530	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.13: Average marginal effects of ANC on child birth—border cities

Panel A: subsamples by birth order					
	No older children		One older child		Two or more older children
Average marginal effect (Quebec×Census1996)	0.0257		0.0084		0.0189
	(0.0084)		(0.0166)		(0.0113)
Number of observations	60,230		21,290		19,780
Panel B: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0245	-0.0070	0.0010	0.0595	0.0047
	(0.0232)	(0.0237)	(0.0205)	(0.0338)	(0.0227)
Number of observations	10,870	10,420	7965	4055	3610
Panel C: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000C–$79,999	C$80,000 higher
Average marginal effect (Quebec×Census1996)	-0.0019	0.0289	0.0305	-0.0021	0.0066
	(0.0123)	(0.0099)	(0.0133)	(0.0268)	(0.0392)
Number of observations	25,610	43,730	23,390	5780	2780
Panel D: subsamples by women's education					
	High school dropout	High school diploma		Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	0.0031	0.0016		0.0260	0.0334
	(0.0189)	(0.0160)		(0.0091)	(0.0131)
Number of observations	15,570	17,710		48,890	19,130
Panel E: subsamples by immigration status					
		Non-immigrant (both parents)		Immigrant (either or both parents)	
Average marginal effect (Quebec×Census1996)		0.0236		0.0054	
		(0.0069)		(0.0164)	
Number of observations		85,770		15,520	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.14: Average marginal effects of ANC on child birth—1986 to 1988 vs1993 to 1995

Panel A: subsamples by birth order					
	No older children		One older child		Two or more older children
Average marginal effect (Quebec×Census1996)	0.0226		0.0135		0.0417
	(0.0038)		(0.0070)		(0.0063)
Number of observations	236,050		78,240		75,740
Panel B: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0156	0.0019	0.0647	0.0609	0.0352
	(0.0111)	(0.0113)	(0.0125)	(0.0178)	(0.0180)
Number of observations	31,385	29,800	29,635	14,875	13,380
Panel C: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Average marginal effect (Quebec×Census1996)	0.0128	0.0263	0.0336	0.0330	0.0122
	(0.0056)	(0.0045)	(0.0064)	(0.0136)	(0.0197)
Number of observations	97,410	168,170	92,300	21,300	10,850
Panel D: subsamples by women's education					
	High school dropout	High school diploma		Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	0.0077	0.0189		0.0285	0.0347
	(0.0070)	(0.0070)		(0.0042)	(0.0071)
Number of observations	72,545	72,775		185,170	59,540
Panel E: subsamples by immigration status					
		Non-immigrant (both parents)		Immigrant (either or both parents)	
Average marginal effect (Quebec×Census1996)		0.0218		0.0248	
		(0.0032)		(0.0088)	
Number of observations		321,250		68,780	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.15: Average marginal effects of ANC on child birth—1984 to 1988 vs1991 to 1995

Panel A: subsamples by birth order					
	No older children		One older child		Two or more older children
Average marginal effect (Quebec×Census1996)	0.0280		0.0018		0.0111
	(0.0037)		(0.0065)		(0.0081)
Number of observations	258,920		71,225		59,880
Panel B: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0009	0.0007	0.0537	0.0589	0.0683
	(0.0126)	(0.0129)	(0.0140)	(0.0203)	(0.0216)
Number of observations	26,070	24,910	18,470	9470	8495
Panel C: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Average marginal effect (Quebec×Census1996)	0.0179	0.0333	0.0386	0.0291	0.0003
	(0.0058)	(0.0046)	(0.0066)	(0.0138)	(0.0196)
Number of observations	97,415	168,165	92,300	21,295	10,845
Panel D: subsamples by women's education					
	High school dropout	High school diploma		Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	0.0148	0.0194		0.0340	0.0427
	(0.0073)	(0.0072)		(0.0043)	(0.0073)
Number of observations	72,545	72,775		185,170	59,540
Panel E: subsamples by immigration status					
		Non-immigrant (both parents)		Immigrant (either or both parents)	
Average marginal effect (Quebec×Census1996)		0.0265		0.0407	
		(0.0033)		(0.0089)	
Number of observations		321,245		68,780	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.16: Average marginal effects of ANC on child birth—excluding immigrants

Panel A: subsamples by birth order					
	No older children		One older child		Two or more older children
Average marginal effect (Quebec×Census1996)	0.0161		0.0149		0.0239
	(0.0041)		(0.0075)		(0.0053)
Number of observations	179,000		67,400		74,840
Panel B: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0233	0.0063	0.0383	0.0432	0.0234
	(0.0105)	(0.0107)	(0.0121)	(0.0186)	(0.0171)
Number of observations	34,515	32,885	28,885	14,470	13,180
Panel C: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Average marginal effect (Quebec×Census1996)	0.0033	0.0136	0.0319	0.0265	-0.0227
	(0.0058)	(0.0045)	(0.0065)	(0.0140)	(0.0210)
Number of observations	79,710	140,000	76,130	17,010	8390
Panel D: subsamples by women's education					
	High school dropout	High school diploma	Some post-secondary	Bachelor degree or higher	
Average marginal effect (Quebec×Census1996)	0.0070	0.0111	0.0192	0.0236	
	(0.0071)	(0.0070)	(0.0043)	(0.0075)	
Number of observations	59,550	60,680	154,090	46,930	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.17: Average marginal effects of ANC on child birth—female aged 25 to 34

Panel A: subsamples by birth order					
	No older children		One older child		Two or more older children
Average marginal effect (Quebec×Census1996)	0.0125		0.0106		0.0285
	(0.0045)		(0.0070)		(0.0050)
Number of observations	162,670		78,860		91,590
Panel B: subsamples by previous children					
	Son	Daughter	Son and daughter	Two sons	Two daughters
Average marginal effect (Quebec×Census1996)	0.0187	0.0024	0.0367	0.0452	0.0358
	(0.0098)	(0.0100)	(0.0106)	(0.0166)	(0.0161)
Number of observations	40,350	38,505	35,140	17,635	15,955
Panel C: subsamples by income group					
	Under C$19,999	C$20,000–C$39,999	C$40,000–C$59,999	C$60,000–C$79,999	C$80,000 higher
Average marginal effect (Quebec×Census1996)	0.0078	0.0116	0.0284	0.0204	-0.0069
	(0.0062)	(0.0046)	(0.0062)	(0.0132)	(0.0194)
Number of observations	72,400	143,310	86,420	20,480	10,510
Panel D: subsamples by women's education					
	High school dropout	High school diploma		Some post-secondary	Bachelor degree or higher
Average marginal effect (Quebec×Census1996)	-0.0003	0.0193		0.0166	0.0260
	(0.0071)	(0.0070)		(0.0044)	(0.0072)
Number of observations	68,680	68,240		167,450	58,140
Panel E: subsamples by immigration status					
	Non-immigrant (both parents)			Immigrant (either or both parents)	
Average marginal effect (Quebec×Census1996)	0.0135			0.0201	
	(0.0033)			(0.0088)	
Number of observations	270,810			62,310	

Notes: Dependent variable is *Had a child*. Robust standard errors are in parenthesis. The number of observations is rounded to the nearest multiple of 5

Table 3.18: Data sources

Source	Title	Reference table or year
Statistics Canada	Induced abortions in hospitals and clinics, annual	CANSIM Table 106-9013
Statistics Canada	Consumer Price Index, annual	CANSIM Table 326-0021
Statistics Canada	Estimates of population, annual	CANSIM Table 051-0001
Statistics Canada	Vital Statistics - Birth Database	Years: 1946 to 2013
Statistics Canada	Census of Population	Years: 1986, 1991, 1996, 2001

Conclusion

This thesis attempts to answer three important questions: 1) Why did India's relative price of investment rise in 80s and fall in 1990s and afterwards? 2) Why is agricultural productivity very low in India? and 3) Did the pro-natalist policy in Quebec accomplish its goal of increasing fertility? Chapter 1 studies the puzzling dynamics of the relative price of capital goods in India between 1980 and 2006, and finds that India's trade policies during that period explains much of the puzzle. Chapter 2 examines why labor productivity in the Indian agricultural sector is low and shows that implementation of a social insurance system in the urban area could have raised labor productivity in the agricultural sector. Chapter 3 studies the effects of a pro-natalist policy in Quebec and finds that Quebec's baby bonus accomplished its goal of increasing fertility.

In Chapter 1, we construct a small open economy model where the government uses tariff and non-tariff barriers to limit the import of foreign capital goods. We calibrate the model to India using data from the Penn World Table and use it to account for the dramatic rise and fall of Indian relative price of investment. Our benchmark calibration implies that the model can generate a 23 percent rise in the relative price of investment between 1981 and 1991 due to increasing distortions created by quantitative restrictions on capital goods imports in the face of a growing economy. The model also accounts for a 28 percent fall in the relative price of investment over the subsequent 15 years as tariff rates fell from 72.7 percent to 7.6 percent and quantity restrictions were removed. The model allows us to separate the impact of tariff rate reductions from the impact of the implicit distortions to investment created by quantity restrictions on capital goods imports. We uncover a considerable general equilibrium impact of these price changes on output and consumption per worker and show that the Indian government's import substitution policies exerted a significant drag on the economy prior to reform. Moreover the removal of capital import restrictions and reduction of tariff rates accounts for one fifth of the observed increase in GDP per worker in India between 1991 and 2006.

In Chapter 2, we develop a tractable quantitative framework by incorporating one potential explanation to address the large labor productivity gap between urban and agricultural sectors in India . If residing in a village provides access to a network that effectively insures against income fluctuations, then households are less willing to live in the cities where labor income risk is uninsured. As a result, labor stays cheap in agriculture, and the incentives for switching to capital-intensive methods of farming remain weak. In order to understand the quantitative importance of this mechanism, we calibrate the model to Indian data and study an abstract policy intervention – a provision of complete insurance against earnings risk in the city. The policy intervention reduces the urban-rural labor productivity gap by 30 percent and raises aggregate labor productivity by 16 percent. This effect comes about because of the 7 percent drop in agricultural share of employment, which encourages an inflow of capital in agricultural sector and raises the average farm size by 12 percent.

Chapter 3 examines the impact of a universal cash transfer policy, the Allowance for Newborn Children (ANC), for all babies born from May 1988 to September 1997 to all residents of Quebec. We find a strong increase in the probability of having a third child or higher order. We are aware that these results are due to the specific payment structure of the ANC. We observe the completed fertility rates of many cohorts that were exposed to the ANC. We see that, in addition to a transitory effect where women were having their children closer together, there was also an increase in completed fertility of women aged 15 to 39, illustrating that the pro-natalist policy does have a permanent effect on fertility in Quebec. We find that among 3-child households the baby bonus was able to create more one daughter-two son families then other sibship sex compositions. Pro-natalist policies, if structured correctly, could cost-effectively increase fertility and alleviate the immense concern of below-replacement rates for developed nations. Furthermore, pro-natalist policies can also diminish gender preferences by incentivizing parents to have more children.