

**AN ANALYSIS OF THE IMPACT OF
GOVERNMENTS' POLICY ON ECONOMIC
ACTIVITIES**

ZIFAN BI

**A DISSERTATION SUBMITTED TO
THE FACULTY OF GRADUATE STUDIES
IN PARTIAL FULFILLMENT OF THE REQUIREMENTS
FOR THE DEGREE OF
DOCTOR OF PHILOSOPHY**

**GRADUATE PROGRAM IN ECONOMICS
YORK UNIVERSITY
TORONTO, ONTARIO**

JULY 2023

© ZIFAN BI, 2023

Abstract

This PhD dissertation explores the intersection of international trade, empirical microeconomics, and public economics. In each chapter of the dissertation, I investigate the impact of free trade agreements (FTAs), preferential tariffs under the Generalized System of Preferences (GSP), and government regulations on local labor markets, respectively.

In the first chapter, the study examines the effects of FTAs on members' incentives to participate in multilateral trade liberalization. By addressing endogeneity issues and considering the parallel existence of multiple FTAs, the research provides robust evidence that FTA formation leads to a reduction in external tariffs towards non-members, shedding light on the role of FTAs in accelerating global free trade.

The second chapter focuses on understanding the factors underlying countries' choices of preferential tariffs towards developing countries under the GSP. A theoretical model is constructed to capture the donor countries' consideration of the economic interests of recipient countries. The findings demonstrate that donors partially internalize the effect of GSP tariffs on profits of firms from developing countries, indicating their concern for the welfare of recipient countries. However, evidence for such considerations among the least developed countries is limited.

In the third chapter, the study analyzes the impact of government regulations on local labor markets, with a specific focus on the United States. Using innovative quantification methods, the research confirms the cost of regulations and their adverse effects on local employment and income equality. Furthermore, the analysis highlights the influence of political party affiliation, showing that regulations under Republican governments tend to attenuate the negative impacts on employment and income equity compared to regulations under Democratic governments.

Dedication

To my beloved parents, wife and daughter, whose unwavering support and belief in me have made all of this possible.

Acknowledgments

I would like to express my deep gratitude and appreciation to the following individuals who have played a crucial role in my academic journey and the successful completion of my PhD dissertation.

First and foremost, I am immensely grateful to my supervisor, Professor Andrey Stoyanov, for his unwavering support, guidance, and expertise throughout the research process. His mentorship has been invaluable in shaping my academic growth and research direction.

I extend my thanks to my committee members, Professor Chaoran Chen and Professor Jun Zhao, for their insightful feedback and contributions not only to my dissertation but also for their assistance and support during my job market preparations.

I am incredibly grateful to Professor George Georgopoulos for believing in my potential and providing me with the opportunity to pursue my PhD. Your guidance and encouragement have been instrumental in my academic and professional development.

I would like to acknowledge the valuable input and suggestions provided by Professor Ben Sand, Professor Ida Ferrara, Professor Antoine Djogbenou, and Professor Nils-Petter Lagerlöf. Their expertise and guidance have significantly enhanced my research quality.

I am grateful for the administrative assistance provided by Regina Pinto, Rex Ferati, Sara Bang, and the late Lisa Guidi. Their support and dedication have been crucial in navigating the administrative aspects of my academic journey.

I would also like to extend my heartfelt appreciation to my cohort members, Anfal Adawi and Nicole Hoseini, as well as my childhood friends Cong Wang and Aijingwen Li. Their friendship, timely assistance, and shared experiences have made this pursuit all the more fulfilling and enjoyable.

Lastly, I would like to extend a special thank you to the entire Department of Economics at York University. The collective knowledge, resources, and intellectual environment provided by the department have significantly contributed to my academic growth.

Thank you all for appearing in my life and making it meaningful.

Contents

Abstract	ii
Dedication	iii
Acknowledgments	iv
Table of Contents	v
List of Tables	vi
Table of Figures	vi
1 Integrated Analysis of the Effect of the Free Trade Agreements on Multi-lateral Trade Liberalization	1
1.1 Introduction	1
1.2 Theoretical Framework	5
1.2.1 The Effect of FTA on MFN Tariffs: Theoretical Background	5
1.2.2 Theoretical Setup	7
1.3 Empirical Implementation	10
1.3.1 Empirical Specification	10
1.3.2 Endogeneity Issues	14
1.3.2.1 Endogeneity from Import Shares	14
1.3.2.2 Endogeneity from Preference Margin	15
1.3.2.3 Endogeneity from Import Decision	16
1.3.2.4 Endogeneity from Parallel FTAs	17
1.3.3 Data	18
1.4 Estimation Results	22

1.4.1	OLS Results	22
1.4.2	IV Results	23
1.4.2.1	IV Results for Instrumenting Import Share	23
1.4.2.2	IV Results for Instrumenting Preference Margin	24
1.4.2.3	IV Results for Instrumenting Import Share and Preference Margin	25
1.4.2.4	IV Results with Parallel FTAs	26
1.5	Replications and Extensions	31
1.5.1	Limão [2006] Replication	31
1.5.2	Development Level	33
1.5.3	Propensity of Joining FTAs	36
1.5.4	Relative Economic Size	38
1.6	Conclusion	38
	Bibliography for Chapter 1	41
	Appendix for Chapter 1	46

2 Did the intended beneficiaries of the Generalized System of Preference (GSP) benefit from it? 51

2.1	Introduction	51
2.2	Theoretical Framework	57
2.2.1	Theoretical Setup	57
2.2.2	Trade policy without Tariff Cooperation	59
2.2.3	Trade policy with Tariff Cooperation	61
2.3	Empirical Implementation	62
2.3.1	Empirical Specification	62
2.3.2	Limitation and Implication of Empirical Model	64
2.3.3	Endogeneity Issues	65
2.3.4	Data	67
2.4	Estimation Results	69

2.5	Robustness	71
2.6	Conclusion	74
	Bibliography for Chapter 2	75
3	The Effect of Regulation on the Employment Rate in the U.S.	79
3.1	Introduction	79
3.2	Empirical Specification	84
3.2.1	Data	84
3.2.2	Empirical Model	86
3.2.3	Model with Disaggregated Regulation	87
3.3	Empirical Results	89
3.4	Conclusion	94
	Bibliography for Chapter 3	96

List of Tables

1.1	Statistics Summary I	19
1.2	Statistics Summary II	20
1.3	OLS Results for the Main Model	22
1.4	IV Results for Instrumenting Import Share	24
1.5	IV Results for Instrumenting Preference Margin	24
1.6	IV Results for Instrumenting Import Share and Preference Margin	25
1.7	IV Results with Parallel FTAs	26
1.8	Comparison of Different Instruments	28
1.9	Replication of Limão [2006]	32
1.10	Replication of Limão [2006] - Our Methodology	32
1.11	Development	35
1.12	FTA Propensity	37
1.13	Relative Economic Size	39
2.1	GSP Donor Country List	67
2.2	Summary Statistics	69
2.3	Estimation Results	72
2.4	Estimation Results of Theoretical Model	73
3.1	Summary Statistics	86
3.2	Estimated Results for the Effect of Regulation on Local Economic Activities	91
3.3	Regulation Added by Different Ruling Parties	94

List of Figures

1.1	IV Results Comparison	30
1.2	IV Result with Confidence Interval	31
1.3	Number of Effective FTAs (1994-2017)	45
1.4	Trade under FTA over Global Trade (1994-2017)	45

Chapter 1

Integrated Analysis of the Effect of the Free Trade Agreements on Multilateral Trade Liberalization

1.1 Introduction

Since the General Agreement on Tariffs and Trade (GATT) was established in 1947, liberalization of the global trade system has proceeded either through unilateral tariff reductions by individual countries, or preferential within a small group of countries, or multilateral between all GATT members. While multilateral trade liberalization has been stalled after the completion of the Uruguay round of WTO negotiations in 1994, many countries switched to preferential liberalization to gain access to markets of their key trade partners, mostly through Free Trade Agreements (FTA). Between 1994 and 2017, the last year of our sample, the number of FTAs has increased almost four-fold (Figure 1.3) and the share of global trade covered by an FTA provision reached 41%, up from 19% in 1994 (Figure 1.4). However, preferential trade liberalization does not guarantee standard welfare gains, unless it is followed by tariff reductions towards non-members large enough for their exports not to decline. Moreover, there are concerns that the possibility to liberalize trade on

bilateral basis may undermine countries' incentives to engage in multilateral liberalization, which further highlights the importance of studying the effect of FTAs on external trade policies of their members.

In this paper we propose a new strategy for exploring the effect of FTA preferential tariff reductions on member countries' most favoured nation (MFN) tariffs. We start by developing a simple partial equilibrium model of optimal MFN tariff before and after FTA formation to motivate our empirical approach. In the model, whether governments respond to preferential trade liberalization with higher or lower MFN tariffs depends on their objectives. If governments maximize national welfare, the standard tariff complementarity effect arises in which a preferential reduction in tariffs towards a group of FTA partners induces them to lower external tariffs. This positive relationship between FTA external and internal tariffs is stronger in industries with large presence of firms from FTA partner countries, because the effectiveness of the MFN tariff to generate revenue and redistribute consumer expenditure towards domestic producers declines by more when the larger share of imports is exempted from tariffs. Conversely, if governments are not pure welfare maximizers but use FTAs as a leverage to extract concessions in non-trade issues, as in [Limão \[2006\]](#), then they may want to maintain higher preference margins and increase gains from the FTA to other members in order to foster their cooperation in non-trade issues. Such considerations work against tariff complementarity, especially in sectors that generate larger surplus for FTA partners.

Using the structure of our model, we proceed to empirically estimating the effect of FTAs on countries' Most Favoured Nation (MFN) tariffs on the sample of all FTAs formed between 1988 and 2017. We find that the tariff complementarity effect of FTAs is persistent and statistically significant, yet the magnitude of the implied effects are much smaller than what was estimated in other studies. Specifically, our results indicate that for an average industry and a pair of FTA partner countries, a one percentage point reduction in FTA tariff leads to 0.012-0.017 percentage points decrease in the MFN tariff. For comparison, similar estimate is 0.2 for Latin American FTAs ([Estevadeordal et al. \[2008\]](#)), and 0.3 for NAFTA ([Mai and Stoyanov \[2015\]](#)). Our mean estimates are relatively small because most

FTAs in our sample are between countries with small bilateral trade flows, while previous studies typically focus on FTAs between main trade partners. Applying our estimates to the Canada-U.S. Free Trade Agreement, we obtain the mean elasticity of the MFN tariffs with respect to preferential tariff in the range of 0.2-0.25, which puts our results in line with much of the previous literature.

As for the effect of cooperation in non-trade issues on FTA tariffs, there is no clear evidence in our data that countries maintain higher preference margin in industries that matter more for their FTA partners. We confirm that the US liberalized its trade policy less in industries where FTA preferences were granted, as in [Limão \[2006\]](#), but that conclusion cannot be generalized to other FTAs. Thus, our analysis indicates that non-trade considerations, which tend to make FTAs more protectionist, only apply to selected economies, such as US and EU, and do not play any major role in an average FTA.

The main challenge that we face in identifying causal effect of preferential tariff reductions on MFN tariffs is that both tariffs can be affected by common economic or political shocks. Although preferential tariff reductions are negotiated prior to FTA commencement and are pre-determined relative to subsequent changes in MFN tariffs, the two can still be determined by persistent economic conditions. Previous studies exploited specific features of FTA they analyzed. Work with a diverse sample of FTAs requires a generalized approach. To break down the simultaneity link between internal and external tariffs, we instrument country's tariff reduction schedule within a given FTA with the corresponding tariff reduction under the first FTA formed by the same country that was established at least five years before. Since governments tend to protect some industries more than other due to fundamental economic and political factors, the tariff reduction schedules of past and current FTA are positively correlated. However, tariff reduction schedules from past FTAs are independent of current shocks and thus represent a valid instrument for current FTA tariff reductions. We show that without addressing endogeneity of preferential tariff reductions, previous studies may have overestimated the effect of FTAs on MFN tariff reduction by a factor of two.

The second identification challenge, that has been generally overlooked in the literature,

is that many countries form multiple FTAs with tariff reduction schedules overlapping in time. As such, the effect of tariff reductions of one FTA on external tariff may pick up the effect of other FTAs which were formed at around the same time. To avoid overestimating the effect of the current FTA, we construct a group of control variables to capture the effect from both the previous and the subsequent agreements signed by the same country.

Much attention has been paid to the effect of trade agreements on tariffs of their members, and the results of this study complement the existing empirical research on this topic. One group of studies found that joining an FTA promotes trade liberalization towards non-members, i.e. that FTAs are “building blocks” to multilateral trade system (Estevadeordal et al. [2008] for Latin American FTAs, Ketterer et al. [2014] and Mai and Stoyanov [2015] for Canada-US FTA). The other group of studies shows that forming an FTA hinders global trade liberalization and act as “stumbling blocks” (Limão [2006] for the US, Karacaovali and Limão [2008] for the EU, Ketterer et al. [2015] for Japan). This difference in results stems from the fact that the two groups investigate different mechanisms through which preferential tariffs affect MFN tariffs. The former group estimates the tariff complementarity effect that induces FTAs to lower external tariffs in response to preferential tariff reduction. In contrast, the latter group tests for the presence of cooperation between member countries in non-trade areas, manifested in higher preference margins in industries that are more important for FTA partners. Finding empirical support for one of the effects without estimating the other does not allow drawing any conclusions on the overall effect of FTAs on member countries’ tariffs. In this paper we estimate both effects and find that for an average FTA only the tariff complementarity effect is empirically relevant. Therefore, our findings support the idea that FTA formation provides member countries with economic incentives to liberalize their trade policies.

Another contribution of our analysis is that we explore relationship between MFN and preferential tariffs based on the sample of all FTAs formed between 1989 and 2017. In contrast, all works mentioned above are case studies focusing on one or a small group of FTAs, and it is unclear whether the results obtained for one specific FTA or country can be generalized to other FTAs. Some authors argue that the difference in the results is driven

by differences in economic development of FTA member countries. Working with a large sample of FTAs allows us to test these hypotheses and identify characteristics of FTAs which are associated with the strength of relationship between external and internal tariffs.

In that respect, our work is closely related to [Kuenzel et al. \[2019\]](#), who also investigate the effect of preferential tariffs on MFN tariffs using a large sample of trade agreements. The authors analyze how MFN tariffs respond to changes in preferential trade shares rather than preferential tariffs. This allows them to bypass some endogeneity problems and the need to instrument preferential tariff changes, which is a challenging task. The intuition behind this approach is that larger FTA preference margins should lead to a bigger increase in the share of imports coming from FTA partners and to stronger tariff complementarity. However, the strength of the tariff complementarity effect depends on the interaction of the preferential tariff and the trade share, and we verify empirically that the empirical model that does not include preferential tariff change is misspecified and can lead to misleading inferences. The results of our estimates, which follow the structure of the theoretical model, imply a substantially weaker tariff complementarity effect.

The rest of the paper is constructed as following. The theoretical framework is discussed in section [1.2](#). Section [1.3](#) introduces the empirical specification (section [1.3.1](#)), endogeneity issues (section [1.3.2](#)) and the data source (section [1.3.3](#)). Section [1.4](#) presents the estimated results for the main model. Section [1.5](#) is the replication for the US and the extensions. Finally, section [1.6](#) concludes with a summary of our findings.

1.2 Theoretical Framework

1.2.1 The Effect of FTA on MFN Tariffs: Theoretical Background

Previous empirical literature on FTAs has extensively explored two key mechanisms through which agreement formation can influence the trade policies of member countries. In this section, we will review and integrate both of these channels into a simplified theoretical

framework of endogenous trade policy, which will serve as a foundation for our empirical model.

The first mechanism we consider is the complementarity of tariffs imposed on different sources of income. When a government lowers tariffs on imports from an FTA partner country, it becomes optimal to reduce tariffs on all other sources of imports as well. There are two main reasons why reducing external tariffs is an optimal response to FTA formation. First, it helps mitigate the distortionary effects of preferential market access on the relative prices of imports from non-FTA countries. By reducing external tariffs, the government can maintain a more level playing field for all imports, regardless of their source. Second, lowering external tariffs aids in recuperating a portion of the tariff revenue loss caused by a shift in consumer expenditure towards imports subject to preferential tariff rates within the FTA. When consumers increase their purchases of goods from FTA partner countries with lower tariff rates, the government experiences a decline in tariff revenue. By reducing external tariffs, the government can partially offset this revenue loss. The positive relationship between external and internal FTA tariffs emerges in any framework with a welfare-maximizing policymaker. We will derive and explain this relationship further using a simple partial equilibrium model.

The second mechanism, proposed by [Limão \[2006\]](#), has received considerable attention in the literature. He argues that FTA tariff preferences and the associated rents can be leveraged by the importing country to induce cooperation from FTA members in non-trade areas. In this scenario, reducing external tariffs upon FTA formation erodes the benefits of the agreement for other members. Consequently, their incentives to cooperate on non-trade issues diminish, and the importing country may choose to maintain higher external tariffs to preserve the leverage provided by the FTA preferences.

By incorporating both mechanisms into our theoretical framework, we aim to provide a comprehensive understanding of the linkages between internal and external tariffs within FTAs.

1.2.2 Theoretical Setup

In the theoretical setup of the model, we consider three groups of countries: a home country indexed by H , a group of countries that have formed FTAs with the home country indexed by $p \in \{1, \dots, P\}$, and all other foreign countries serving the home country's market under the Most Favored Nation (MFN) rule, indexed by $f \in \{1, \dots, F\}$. In total, there are $(1 + P + F)$ countries indexed by c .

All countries have the ability to produce and trade goods in $(1 + I)$ industries. The first good is treated as the numeraire, which is freely traded and produced with labor using constant returns to scale technology in a perfectly competitive market. As a result, the wage is fixed and normalized to a value of unity. For the remaining $i \in \{1, \dots, I\}$ industries, firms from the three groups of countries compete in monopolistically competitive markets with differentiated products. In each industry i , there are n_{ic} firms located in country c , with each firm producing a variety of product i at country-industry-specific marginal costs denoted by c_{ic} . Consequently, the global market consists of $n_{iH} + \sum_{p=1}^P n_{ip} + \sum_{f=1}^F n_{if}$ varieties of product i traded.

The representative consumer at home is assumed to have a quasi-linear Cobb-Douglas utility function:

$$U = X_0 + \sum_i^I a_i \ln X_i, \quad \sum_i^I a_i = 1, \quad (1.1)$$

where a_i is the share of total expenditure that home consumer spends on product i , and X_0 is consumption of the numeraire product. X_i is the constant elasticity of substitution (CES) aggregate consumption of product i derived from the consumption of different varieties of the product:

$$X_i = \left(n_{iH} d_{iH}^{\frac{1}{\sigma_i}} x_{iH}^{\frac{\sigma_i-1}{\sigma_i}} + \sum_{p=1}^P n_{ip} d_{ip}^{\frac{1}{\sigma_i}} x_{ip}^{\frac{\sigma_i-1}{\sigma_i}} + \sum_{f=1}^F n_{if} d_{if}^{\frac{1}{\sigma_i}} x_{if}^{\frac{\sigma_i-1}{\sigma_i}} \right)^{\frac{\sigma_i}{\sigma_i-1}}, \quad (1.2)$$

where d_{ic} is the parameter that captures consumer's preference for product i produced in country c , x_{ic} represents consumption of individual variety of product i originating in

country c , and $\sigma_i > 1$ is the elasticity of substitution between varieties of product i .

Maximizing utility function (1.1) subject to the standard budget constraint, where income is normalized to 1, allows us to derive the demand function for each variety of product i :

$$x_{ic} = a_i d_{ic} p_{ic}^{-\sigma_i} P_i^{\sigma_i - 1} \quad \forall c \in \{H, P, F\}, \quad (1.3)$$

where P_i is the corresponding CES price index for product i

$$P_i = \left(n_{iH} d_{iH} p_{iH}^{1-\sigma_i} + \sum_{p=1}^P n_{ip} d_{ip} p_{ip}^{1-\sigma_i} + \sum_{f=1}^F n_{if} d_{if} p_{if}^{1-\sigma_i} \right)^{\frac{1}{1-\sigma_i}}, \quad (1.4)$$

and p_{ic} is the consumer price at Home of a variety of product i imported from country c .

Next, let us consider the scenario when the home country government imposes ad-valorem tariffs τ_{ip} and τ_{iF} on imports from FTA partner countries and other foreign countries, respectively. Therefore, the profit of a firm from country c selling product i at Home is

$$\pi_{ic} = \left(\frac{p_{ic}}{1 + \tau_{ic}} - c_{ic} \right) q_{ic} \quad \forall c \in \{H, P, F\}, \quad (1.5)$$

where q_{ic} is the quantity of product i supplied to the home country's market, and $\frac{p_{ic}}{1 + \tau_{ic}}$ is the producer price of product i received by the firm located in country c , with $\tau_{iH} = 0$.

Assuming that firms take the aggregate price index as given, the profit-maximizing pricing strategy is

$$p_{ic} = \left(\frac{\sigma_i}{\sigma_i - 1} \right) c_{ic} (1 + \tau_{ic}) \quad \forall c \in \{H, P, F\}, \quad (1.6)$$

and the equilibrium profit function becomes

$$\pi_{ic} = - \frac{p_{ic} q_{ic}}{\sigma_i (1 + \tau_{ic})} \quad \forall c \in \{H, P, F\}. \quad (1.7)$$

Next, we examine the optimal trade policy pursued by the home country government.

At any point in time, the government is assumed to choose a vector of MFN tariffs τ_F that maximizes the sum of consumer surplus, producer surplus and tariff revenue, taking FTA tariffs as given. We treat FTA preferential tariffs as pre-determined because they are negotiated before the agreement comes into force and cannot be adjusted once it is in place.

The government's objective function is then expressed as follows:

$$G(\tau_F) = CS(\tau_F) + PS(\tau_F) + TR(\tau_F) \quad (1.8)$$

where consumer and producer surpluses are given by

$$CS(\tau_F) = U(X_0, X_i, \tau_F) - \sum_{i=1}^I \sum_{c=1}^{H+P+F} n_{ic} p_{ic} x_{ic}, \quad (1.9)$$

$$PS(\tau_F) = \sum_{i=1}^I n_{iH} \pi_{iH} \quad (1.10)$$

and the tariff revenue is equal to

$$TR(\tau_F) = \sum_{i=1}^I \left(\sum_{p=1}^P \frac{\tau_{ip}}{1 + \tau_{ip}} p_{ip} x_{ip} + \frac{\tau_{iF}}{1 + \tau_{iF}} \sum_{f=1}^F p_{if} x_{if} \right). \quad (1.11)$$

Given the partial equilibrium setting, the equilibrium tariff on product i is independent of tariffs and prices in other industries, and we can focus on one industry at a time. The optimal MFN tariff of the Home country on imported good i that maximizes objective function $G(\tau_F)$ is given by:¹

$$\varepsilon_i \frac{\tau_{iF}}{1 + \tau_{iF}} = \frac{\sigma_i - 1}{\sigma_i} s_{iH} + (\sigma_i - 1) \sum_{p=1}^P s_{ip} \frac{\tau_{ip}}{1 + \tau_{ip}}, \quad (1.12)$$

where $s_{iH} = \frac{p_{iH} x_{iH}}{a_i}$ is the share of firms from Home country in the domestic market, defined as the ratio of expenditure that consumers at Home spend on domestic varieties of product i to total consumer expenditure on product i . The term s_{ip} is the share of firms from a partner country p in Home's market, and $\varepsilon_i = (\sigma_i - 1) s_{iF} - \sigma_i$ is the import demand

¹Derivations of optimal trade policies under various trade regimes are presented in the Appendix.

elasticity for product i imported from all non-agreement foreign countries, where s_{iF} is defined similarly to s_{iH} and s_{ip} .

The above expression shows that when a country has preferential trade agreements with a group of other countries, its welfare-maximizing tariff for imports from all other countries is determined by two factors. First, tariffs tend to be larger in industries with greater presence of domestic firms in the market, as reflected by the first term on the right-hand side of equation (1.12). This effect stems from imperfectly competitive structure of the market, which stimulates policymakers to shift consumer expenditure towards domestic varieties of the product and increase profits of domestic firms.

The second term in on the right-hand side of (1.12) is the tariff complementarity effect, identified in previous studies, such as [Bagwell and Staiger \[1999\]](#). It demonstrates positive relationship between tariffs on two different sources of imports. Intuitively, a reduction in tariffs on imports from FTA countries leads to a reduction in imports from the rest of the world. An optimal response of a welfare-maximizing government to FTA formation is to lower tariff rate on imports from non-FTA countries in order to increase imports from countries that are subject to higher rates. Importantly, tariff complementarity effect is stronger when the market share of partner country firms is greater, because the effect of tariff preferences on reduction in imports from non-FTA countries is stronger when s_{ip} is larger.

1.3 Empirical Implementation

1.3.1 Empirical Specification

As mentioned in the previous section, the main theoretical basis of this project is mainly stemmed from the aforementioned papers and we summarize and combine the two previous channels to estimate the general effect of FTAs on the global trade liberalization. However, our empirical specification differentiates from the established research approaches in four ways.

First, in contrast to most of the previous literature, which only estimated one FTA or several FTAs of one country, the construction of the specification model is difficult in this paper since almost all FTAs in the world are considered. One problem needs to be noticed is that a country may have signed more than one FTA and multiple FTAs could be simultaneously effective in a certain period. Therefore, the negotiated preferential tariffs towards different partner countries under different FTAs are different and changing in various patterns. [Estevadeordal et al. \[2008\]](#) bypassed this problem by taking the minimum preferential tariff among all partners as the fixed preferential tariff of the importing country. However, this method has limitation in estimating the effect of all FTAs in the global case, because different FTAs formed by different countries, which are identified by various development levels or other factors, may trigger different effects on MTL in terms of both magnitude and direction. Hence, it is not rational to aggregate the preferential tariffs within one country in this paper, and each FTA would better be treated separately.

Second, our theoretical model (1.12) incorporates the partner-occupied market share as a weighting factor for preferential tariffs. This approach is justified because if the “tariff complementarity” effect exists, it is expected to be stronger when partner countries have a larger share of the home market for a specific product. The underlying reasoning is that when partner countries hold a significant market share in the home country, a reduction in tariffs on imports from these countries is likely to have a more pronounced impact on trade flows and consumer preferences. Consequently, the preferential tariff treatment afforded to these partner countries would exert a greater influence in diverting imports away from non-preferential trade agreement countries.

Third, we have relaxed the assumption that only one effect, either the “building block” effect or the “stumbling block” effect, exists within a country’s FTAs. Instead, we allow for the coexistence of both effects simultaneously. Consequently, based on our theoretical model (1.12), we construct a dummy variable similar to the approach used by [Limão \[2006\]](#). This dummy variable allows us to compare the changes in MFN tariffs for goods covered by FTAs versus goods not covered by FTAs. The aim is to investigate whether the home country maintains higher MFN tariffs in certain “important” industries (FTA goods) for

non-trade benefits.

Finally, to address the issue of unobserved time-invariant heterogeneity, we employ the first-difference approach, as commonly done in previous literature. However, it is important to note that the “building block” and “stumbling block” effects of each FTA on a country’s MFN tariffs accumulate over time. Thus, examining the varying degree of difference between MFN tariffs in two successive years is insufficient to capture the cumulative effect. To account for the cumulative impact of FTAs, our approach involves constructing eight regression equations to estimate the cumulative effect for each individual FTA. This estimation considers the existence period of each FTA, ranging from one year up to eight years. By doing so, we can capture the dynamic nature of the “building block” and “stumbling block” effects over time and obtain a comprehensive understanding of the cumulative impact of FTAs on a country’s MFN tariffs.

Our empirical specification is primarily motivated by our theoretical model (1.12). However, certain modifications are necessary for practical reasons and due to data limitations. Firstly, in model (1.12), the first term (s_{iH}) on the right hand side representing the domestic home market share will be removed from our empirical specification model due to several reasons. First, the available trade and tariff data are at the 6-digit HS product classification level, while the necessary output data for constructing the domestic market share is only available at the 4-digit ISIC level for some developed countries. The aggregation process required would result in a substantial loss of over 80 percent of the observations. Second, Mai and Stoyanov [2015] include the domestic home market share in analyzing the effect of Canada-United States Free Trade Agreement on the Canadian MTL, they also show that the domestic market share has no statistically significant impact on the Canadian MFN tariff.

Secondly, we do not include the elasticity of substitution (σ_i) and the import demand elasticity (ε_i) in our specification model. This is due to the lack of reliable and accurate measurement of these elasticity terms. Including these terms would introduce noise and uncertainty into the regression analysis.

Taking these factors into account, our final specification model is as follows:

$$\Delta_t MFN_{cin} = \beta_0 + \beta_1 \Delta_t PREF_{cin} * Pshare_{cin} + \beta_2 G_{cin} + \alpha_c + \alpha_i + \alpha_n + \varepsilon_{cin}. \quad (1.13)$$

In the above equation, the subscripts t , c , i , and n represent the indices for the year gap, country, industry, and number of FTAs, respectively. The dependent variable, $\Delta_t MFN_{cin}$, is constructed as the t -year MFN tariff margin ($\Delta_{-1} MFN_{cin} - \Delta_t MFN_{cin}$). Specifically, $\Delta_{-1} MFN_{cin}$ is the MFN tariff that country c imposes on industry i in the year before the formation of the n^{th} FTA that country c formed. $\Delta_t MFN_{cin}$ is similar as $\Delta_{-1} MFN_{cin}$ except it is the MFN in the t^{th} year after the formation of the n^{th} FTA that country c formed.

The first explanatory variable, $\Delta_t PREF_{cin}$, is the t -year preferential tariff margin, which is calculated as the difference between the MFN tariff in the year prior to the formation of the n^{th} FTA ($\Delta_{-1} MFN_{cin}$) and the preferential tariff imposed by country c on industry i in the t^{th} year after the formation of the n^{th} FTA ($\Delta_t PREF_{cin}$). The variable $Pshare_{cin}$ represents the home market share occupied by the partners of country c in industry i . This market share is fixed in the year prior to the formation of the n^{th} FTA of country c . We anticipate that, on average, all three variables mentioned above ($\Delta_t MFN_{cin}$, $\Delta_t PREF_{cin}$, and $Pshare_{cin}$) will have positive values. This expectation aligns with the global trend of declining tariffs and the non-negative nature of market share.

The coefficient β_1 captures the “building block” effect, indicating the relationship between changes in internal and external tariffs. We anticipate the coefficient to be positive and statistically significant, confirming the existence of the complementarity effect between the changes in internal and external tariffs. Moreover, the complementarity effect is stronger as the partner-occupied home market share is increasing.

The second explanatory variable, G_{cin} , a dummy variable based on the study by [Limão \[2006\]](#). It takes a value of one if country c imported goods in industry i from any partners under its n^{th} FTA in the year before the formation of that FTA, and zero otherwise. Intuitively, G_{cin} indicates that whether or not an industry is considered “important” to the

FTA partners.

The coefficient β_2 captures the “stumbling block”. If this coefficient is negative and statistically significant, it supports the finding from [Limão \[2006\]](#), indicating that the home country keeps higher MFN tariffs in certain “important” industries (FTA goods) for non-trade benefits.

The fixed effects α_c , α_i , and α_n control for the unobserved country, industry, and FTA characteristics, respectively. The error term is denoted by ε_{cin} . One thing we need to pay attention is that “ Δ_t ” is not an actual dimension in the above regression. It is used to identify the cumulative effect span. For example, when $t = 8$, we estimate the 8-year cumulative effect of FTAs on MTL. This is calculated as the MFN tariff in the year before the formation of the current FTA minus the MFN tariff in the 8th year after the formation of the current FTA. This long difference method captures the tariff margins for each FTA and controls for time-invariant characteristics. In this paper, we analyze cumulative effects up to 8 years, so $1 \leq t \leq 8$.

Regarding the sample construction, all FTAs formed before 1988 are excluded since they fall outside the examination period of this study. The first FTA of each country after 1988 is used to construct instrumental variables for later use. Additionally, all FTAs formed within 5 years of the formation of the first FTA after 1988 for each country are removed to enhance the exogeneity of the earlier instrumental variables. The instruments will be further discussed in the next section.

1.3.2 Endogeneity Issues

To address potential endogeneity issues in our empirical specification, we carefully analyze various sources of endogeneity.

1.3.2.1 Endogeneity from Import Shares

The first concern is that changes in the home country’s MFN tariff may influence the import share of partner countries. While we fixed the market share in the year before the formation of the FTA to enhance the validity of this variable, there is a negotiation

period before FTA formation where potential partners anticipate the adjustment of the home country’s MFN tariffs in the following year. To address this problem, we use the instrumental variable proposed by [Kuenzel et al. \[2019\]](#). This instrument, known as the “export share”, is constructed as the total export value of partner countries relative to the global trade value of the corresponding product (excluding exports to the home country). It is highly effective because it captures the comparative advantage of partner countries in a given industry and strongly correlates with the import share of the home country. Importantly, the “export share” should not impact the home country’s MFN tariff.

1.3.2.2 Endogeneity from Preference Margin

The second endogeneity issue arises from the preference margin variable. Changes in a country’s trade policy may affect its decision to form FTAs, creating reverse causation between changes in MFN tariffs and preferential tariff adjustments. Additionally, international economic shocks or domestic political factors may simultaneously influence the adjustment of MFN and preferential tariff rates, undermining the causality between the preference margin and the MFN margin. To address this, we need to construct an appropriate instrumental variable. In fact, due to the data complexity and its global-scale nature, the selection of the instrument in this case is challenging even though previous literature has made a few attempts. One common example is to use the one-year lagged preferential tariff since it seems to be predetermined relative to the adjustment on the external tariff. However, it may not be valid since the adjustment of the external tariff is highly likely to be anticipated. Another prominent attempt from [Estevadeordal et al. \[2008\]](#) is to use the reciprocal preferential tariff cut from the other FTA partners to instrument the preferential tariff from the home country. This instrument may also lose a part of its validity in estimating the effect of all FTAs on the global trade liberalization since countries under the same FTA may have dramatic differences in economic size or political environment. Therefore, the correlation between the preferential tariffs from different FTA partners is weak especially in the same industry for the two following reasons. First, this reciprocity cannot be guaranteed or at least not in trade realm for the global case. Second, if one

country is protecting the profits of domestic producers from one industry, it may make some tariff concessions for some other industries.

This paper is trying to contribute in constructing the appropriate instrumental variable in the global case. As we mentioned in the previous section, we take the corresponding preferential tariff margin during the country's first FTA as the instrument for the preference margin during the country's other FTAs. For example, in the case of a three-year preference margin, where t equals to 3, the instrumental variable for the explanatory variable $\Delta_3 PPEF_{cin}$ is constructed as the preferential tariff in the year before the formation of country c 's first FTA minus the preferential tariff in the 3^{rd} year after the formation of the first FTA in industry i , denoted by $\Delta_3 PPEF_{cin=1}$. The instrumental variable and the explanatory variable are highly correlated since the protectionist in industry i of country c is barely varying as time and towards different trade partners. Furthermore, the instrumental variable tends to be uncorrelated with the dependent variable and the potential political and economic factors affecting both internal and external tariffs, since all the FTAs of a country in the sample is formed at least 5 years apart from the first FTA of the same country.

1.3.2.3 Endogeneity from Import Decision

The third potential endogeneity issue stems from the dummy variable G , which the basic idea originates from [Limão \[2006\]](#). Intuitively, this variable is an indicator of whether the partner countries has exported to the home country for a given industry. However, one may argue the existence of the reverse causation that the trade policy in the importing country may conversely affect the trade decision of the exporting country. [Limão \[2006\]](#) constructed a dummy variable indicating if the partner countries exported to the US before the Uruguay Round, and this dummy works as the instrument for the G variable from his paper. It is reasonable since [Limão \[2006\]](#) only estimates the change in the MFN tariff of the US before and after the Uruguay Round. Under the circumstances in this paper, however, we estimate the effect of each FTA on the global trade liberalization by examining a longer and smoother test period. As a result, we have to fix the explanatory variable G as the

trade decision in the year before the formation of the FTA, otherwise, it may arise extra time-varying issues for the variable G . Nevertheless, it still may not enough to support exogeneity for the variable G since, again, the exporting country may have the expectation on the adjustment of the importing country's MFN tariff.

To solve this reverse causality problem, we construct another dummy variable, G^{iv} , which takes the value of one if the partners export goods in industry i to the rest of the world except the home country in the year before the formation of the FTA, and zero otherwise. This is a valid instrument since, on one hand, if the exporting country exports products in an industry to other countries shows that the exporting country has the comparative advantage in this industry to a certain extent, so it is highly likely that the partners also export the products to the FTA partners. On the other hand, whether the partners decide to export to other countries is irrelevant to the home country's own trade policy.

1.3.2.4 Endogeneity from Parallel FTAs

The last possible endogeneity case emanates from the parallel FTAs. As we mentioned in section (1.3.1), the home country may have signed more than one FTAs and it is highly possible that the periods of the multiple FTAs may be overlapped with each other. For example, the fifth year of the 1st FTA could be the first year of the 2nd FTA if the 2nd FTA is signed exactly five years after the 1st FTA by the same home country. This arises the problem that we may over estimate the effect of the current FTA if the next FTA is signed within 8 eight years after the formation of the current FTA. Therefore, in order to control the effects of the subsequent overlapped FTAs, we update the specification (1.13) as the following:

$$\begin{aligned} \Delta_t MFN_{cin} = & \beta_0 + \beta_1 \Delta_t PREF_{cin} * Pshare_{cin} + \beta_2 G_{cin} + \\ & \sum_{k=1}^{t-1} \Phi_k \Delta_{t-k} PREF_{cin+k} * Pshare_{cin+k} + \alpha_c + \alpha_i + \alpha_n + \varepsilon_{cin}. \end{aligned} \quad (1.14)$$

The above specification (1.14) is the same as specification (1.13), except it includes

an extra summation term $\sum_{k=1}^{t-1} \Phi_k \Delta_{t-k} PREF_{cin+k} * Pshare_{cin+k}$. This is the summation of the weighted preferential tariffs from all the later FTAs that are overlapped with the current FTA. Since this project is estimating the cumulative effect of FTAs from 1-year to 8-year ($1 \leq t \leq 8$), the number of the weighted preferential tariffs from the subsequent FTAs in the summation term is varying. For example, when $t = 8$, the summation term would include 7 weighted preferential tariffs from 7 subsequent FTAs in maximum if the current FTA and all the subsequent FTAs are formed in successive years. When $t = 1$, the summation term includes zero term, and the specification (1.14) is exactly the same as the specification (1.13). The coefficient Φ_k is capturing the “building block” effect of the $(n + k)^{th}$ FTA.

Overall, by addressing these potential endogeneity issues through appropriate instrumental variables and specification adjustments, we aim to enhance the validity and robustness of our estimations.

1.3.3 Data

In this paper, we collect data from various sources, focusing on the period between 1988 and 2017. The sample consists of 46 countries and 5413 industries. The information on Free Trade Agreements (FTAs) is obtained from the RTA database of the World Trade Organization. Trade data, including MFN and preferential tariff data, is sourced from the World Integrated Trade Solution (WITS) and is available at the 6-digit HS product classification level. The concordance of country names and numeric codes is also provided by WITS. National income levels data is sourced from The World Bank. Gross Domestic Product (GDP) data is collected from the World Development Indicators.

In cases where a home country has formed an FTA with multiple partners, we weight the preferential tariffs towards different partners based on their respective export shares to the home country. The partner-occupied home market share ($Pshare_{cin}$) theoretically represents the imports from FTA partner countries relative to domestic consumption. However, due to the limitations of the available 6-digit HS industry-level output data, we are unable to construct a perfect measure of domestic consumption. As a proxy, we use

the corresponding import share (imports from partners under the n^{th} FTA relative to total imports) to approximate the home market share supplied by FTA partners.

Note that we treat the European Union (EU) as a single entity for three reasons. First, the EU engages in trade with non-members as an integrated economic entity. Second, the EU sets common external tariffs, which can be considered as MFN tariffs towards non-members. Third, despite changes in the number of EU members during the examining period, including four enlargements, no member has withdrawn from the EU.

Table 1.1: Statistics Summary I

Variable	Mean	St. Dev.	Min	Max	Observations
$\Delta_1 MFN$	0.24	2.27	-85.67	30	193,532
$\Delta_1 PREF$	1.59	4.32	-84.5	73	193,532
$\Delta_1 PREF^{iv}$	2.29	4.55	-88.83	61.25	193,532
$\Delta_2 MFN$	0.34	2.65	-85.67	100	197,391
$\Delta_2 PREF$	2.42	5.32	-84.5	100	197,391
$\Delta_2 PREF^{iv}$	3.02	5.31	-89.76	90	197,391
$\Delta_3 MFN$	0.32	2.83	-100	100	202,894
$\Delta_3 PREF$	2.39	5.03	-84.5	100	202,894
$\Delta_3 PREF^{iv}$	3.64	6.74	-80.06	100	202,894
$\Delta_4 MFN$	0.73	3.88	-97	65	190,803
$\Delta_4 PREF$	3.18	6.63	-84.5	92.5	190,803
$\Delta_4 PREF^{iv}$	4.54	8.62	-81	97	190,803
$\Delta_5 MFN$	0.56	3.53	-86	79.43	187,406
$\Delta_5 PREF$	2.24	5.08	-28.08	87.81	187,406
$\Delta_5 PREF^{iv}$	4.03	8.06	-83.33	83.75	187,406
$\Delta_6 MFN$	0.55	3.51	-85.67	79.43	186,704
$\Delta_6 PREF$	2.53	5.43	-40	87.81	186,704
$\Delta_6 PREF^{iv}$	4.36	8.38	-86.67	97	186,704
$\Delta_7 MFN$	0.64	3.96	-86	90.83	175,614
$\Delta_7 PREF$	3.10	6.67	-24	100	175,614
$\Delta_7 PREF^{iv}$	4.80	8.45	-75.83	100	175,614
$\Delta_8 MFN$	0.46	2.83	-85.67	90.83	191,372
$\Delta_8 PREF$	3.02	6.37	-23.5	100	191,372
$\Delta_8 PREF^{iv}$	4.82	7.83	-79.33	100	191,372

Table 1.1 provides the statistics summary for the tariff margin variables in this project. $\Delta_t MFN$ is the dependent variable representing the average cumulative MFN tariff reduction over a t -year period. It can be observed that the average cumulative MFN reduction increases

Table 1.2: Statistics Summary II

Variable	Mean	St. Dev.	Variable	Mean	St. Dev.	Observations
$G_{t=1}$	0.30	0.46	$Pshare_{t=1}$	0.04	0.14	193,532
$G_{t=1}^{iv}$	0.46	0.50	$Pshare_{t=1}^{iv}$	0.03	0.08	193,532
$G_{t=2}$	0.28	0.45	$Pshare_{t=2}$	0.04	0.14	197,391
$G_{t=2}^{iv}$	0.42	0.49	$Pshare_{t=2}^{iv}$	0.02	0.08	197,391
$G_{t=3}$	0.31	0.46	$Pshare_{t=3}$	0.05	0.16	202,894
$G_{t=3}^{iv}$	0.48	0.50	$Pshare_{t=3}^{iv}$	0.03	0.09	202,894
$G_{t=4}$	0.28	0.45	$Pshare_{t=4}$	0.04	0.15	190,803
$G_{t=4}^{iv}$	0.42	0.49	$Pshare_{t=4}^{iv}$	0.03	0.08	190,803
$G_{t=5}$	0.30	0.46	$Pshare_{t=5}$	0.05	0.16	187,406
$G_{t=5}^{iv}$	0.44	0.50	$Pshare_{t=5}^{iv}$	0.03	0.09	187,406
$G_{t=6}$	0.31	0.46	$Pshare_{t=6}$	0.05	0.17	186,704
$G_{t=6}^{iv}$	0.44	0.50	$Pshare_{t=6}^{iv}$	0.03	0.09	186,704
$G_{t=7}$	0.29	0.45	$Pshare_{t=7}$	0.04	0.15	175,614
$G_{t=7}^{iv}$	0.41	0.49	$Pshare_{t=7}^{iv}$	0.03	0.09	175,614
$G_{t=8}$	0.33	0.47	$Pshare_{t=8}$	0.05	0.16	191,372
$G_{t=8}^{iv}$	0.50	0.50	$Pshare_{t=8}^{iv}$	0.03	0.09	191,372

as the time period extends. For example, the one-year cumulative reduction is 0.24%, while the eight-year cumulative reduction is 0.46%. The variable $\Delta_t PREF$ represents the average cumulative reduction in preferential tariffs over a t -year period. It is expected to be larger than the corresponding MFN margin, and the data confirms this expectation. The average cumulative preferential tariff reduction ranges from 1.59% for the one-year cumulative reduction to 3.02% for the eight-year cumulative reduction. The variable $\Delta_t PREF^{iv}$ serves as the instrumental variable for $\Delta_t PREF$. It is constructed similarly to $\Delta_t PREF$, but it represents the preferential tariff reduction under the first FTA formed by the same home country. Intuitively, the instrumental variable is expected to be greater than $\Delta_t PREF$ since there is more room for tariff variation under the earlier FTAs. The average cumulative preferential tariff reduction for $\Delta_t PREF^{iv}$ increases from 2.29% for the one-year cumulative cut to 4.82% for the eight-year cumulative cut.

All three variables, $\Delta_t MFN$, $\Delta_t PREF$, and $\Delta_t PREF^{iv}$, capture changes in tariff rates. It is noteworthy that these variables exhibit an increasing pattern, which aligns with the global trends of changing tariffs over time.

Table 1.2 presents the statistics summary for variables ranged from 0 to 1. The variable G is the explanatory variable capturing the “stumbling block” effect, and G^{iv} is the instrumental variable for G . We have already eliminated the FTAs that had formed before 1988 and after 2010 to make sure that we are able to trace the first to the eighth year of each FTA in our sample. Therefore, technically, we expect G and G^{iv} to be stable across t since they are fixed in the year before the FTA. However, due to the availability of tariff data and sample restrictions, the number of observations may vary slightly across the eight regressions. To calculate the average “stumbling block” effect in each case, we use $G_{t=t} (G_{t=t}^{iv})$ to represent the average value of $G(G^{iv})$ in each sample. On average, the values of G range around 0.30, indicating that approximately 30% of industries were already traded between the home country and its FTA partners even before the formation of the FTA. This suggests the existence of pre-existing trade relations between the parties. Alternatively, it can be explained as, on average, 30% of industries are “important” to the FTA partners. The values of G^{iv} average around 0.45, indicating that, on average, the FTA partners exported around 45% of industries to the rest of the world, excluding the home country, before the formation of the FTA. This highlights the economic significance of FTA partners in global trade. The variable $Pshare$ represents the average import share from FTA partners in the year before the formation of each FTA. Similar to G , $Pshare$ varies across different regression samples due to data availability. In the first year of each FTA, the average import share from FTA partners, denoted as $Pshare_{t=1}$, is 0.04, indicating that 4% of the total import value comes from FTA partners. The instrumental variable $Pshare_1^{iv}$ has a value of 0.03, indicating that in the first year of each FTA, the average share of FTA partners’ exports to the rest of the world, relative to the global trade value of a certain product, is 3%. This instrumental variable is used to address the potential endogeneity issue related to the variable $Pshare$.

Overall, these statistics from table 1.2 provide insights into the prevalence of trade relations, the significance of FTA partners in global trade, and the characteristics of import shares before the formation of FTAs.

1.4 Estimation Results

1.4.1 OLS Results

Table 1.3: OLS Results for the Main Model

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t PREF$	0.1402*	0.1790*	0.0552	0.2101***	0.3782**	0.2253***	0.1836***	0.1099**
$\times Pshare$	(1.75)	(1.96)	(1.59)	(3.40)	(2.62)	(3.87)	(3.39)	(2.36)
G	-0.0204 (-0.48)	-0.0778 (-1.34)	0.0380 (0.66)	-0.0315 (-0.36)	-0.1245 (-1.32)	-0.1630 (-1.64)	-0.0950** (-2.27)	0.0041 (0.04)
$Adj.R^2$	0.2424	0.1760	0.2055	0.1766	0.2557	0.2228	0.2723	0.2911
N	193,532	197,391	202,894	190,803	187,406	186,704	175,614	191,372

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

In this section, we present the OLS results for the model specified in section 1.3.1. Table 1.3 displays the OLS results for specification (1.13). To control for a wide range of unobserved heterogeneity, all regressions include country, industry, and FTA year fixed effects. The standard errors are double clustered at the country and industry levels.

The dependent variables in columns (1) to (8) are the cumulative MFN tariff reductions ($\Delta_t MFN$) for 1-year to 8-year periods, respectively. The first explanatory variable, $\Delta_t PREF * Pshare$, represents the interaction between the cumulative preferential tariff reduction and the import share of FTA partners. The estimated coefficients β_1 for the variable $\Delta_t PREF Pshare$ are 0.14, 0.18, 0.06, 0.21, 0.38, 0.23, 0.18, and 0.11 for t equal to 1 to 8, from column (1) to column (8). Except for the insignificant coefficient of 0.06 (significant at 15% confidence level) in column (3), all other coefficients are positive and significant as expected. This indicates that the cumulative reduction in preferential tariffs, weighted by the import share of FTA partners, has a positive and significant impact on the cumulative MFN tariff reductions.

The second explanatory variable, G , is derived from [Limão \[2006\]](#) and is used to identify if an industry is considered "important" to the FTA partners. The estimated coefficients β_2 for variable G are -0.02, -0.08, 0.04, -0.03, -0.12, -0.16, -0.10, and 0.00 in columns (1) to (8), respectively. Except for the coefficients of 0.04 and 0.00 in columns (1) and (8), which are positive, the estimates of β_2 in the remaining columns are negative. However, only the coefficient of -0.10 in column (7) is significant at the 5% confidence level. The lack of statistical significance for the estimated coefficient of G , suggests that there is limited evidence to support the presence of a consistent "stumbling block" effect in trade liberalization across the examined FTA cases.

As discussed in section [1.3.2](#), considering the possibility of endogeneity issues, we will refrain from providing an economic interpretation of the magnitude of the estimated coefficients in this section. The next section will delve further into the analysis.

1.4.2 IV Results

In this section, we will present the IV results for specification [\(1.13\)](#) and [\(1.14\)](#) by instrumenting the explanatory variable $\Delta_t PEF * Pshare$ in different ways. By comparing and analyzing these sets of IV results, we aim to identify the most effective approach to address the endogeneity issues in our analysis.

1.4.2.1 IV Results for Instrumenting Import Share

As discussed in section [1.3.2](#), the first endogeneity problem is that the home country's trade policy may in turn affect the partner country's trade volume. Table [1.4](#) presents the second-stage IV results for specification [\(1.13\)](#), when we instrument the explanatory variable $\Delta_t PEF * Pshare$ using the instrumental variable $\Delta_t PEF * Pshare^{iv}$. This instrumental variable means that we only instrument the "weighted share" ($Pshare$) of the weighted preferential tariff cut ($\Delta_t PEF * Pshare$) by the corresponding export shares to the rest of the world ($Pshare^{iv}$), of which the idea comes from [Kuenzel et al. \[2019\]](#).

Table 1.4: IV Results for Instrumenting Import Share

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t PREF$ $\times Pshare$	0.5046 (1.24)	0.3752*** (2.72)	0.2104*** (2.97)	0.4932* (1.96)	0.5509** (2.24)	0.3515*** (3.50)	0.3666*** (3.67)	0.4041 (1.38)
G	-0.0524 (-0.84)	-0.0809 (-1.13)	0.0879 (1.00)	-0.0450 (-0.29)	-0.2822 (-1.45)	-0.2920* (-1.70)	-0.1394 (-1.41)	0.0258 (0.20)
N	193,532	197,391	202,894	190,803	187,406	186,704	175,614	191,372

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

The IV estimates for β_1 equal to 0.50, 0.38, 0.21, 0.49, 0.55, 0.35, 0.37, 0.40 from column (1) to column (8). Except 0.50 and 0.40 from column (1) and column (8) are insignificant, the values from the rest columns are all positive and significant. We also instrument the “important” industry indicator G using another indicator that if the FTA partner counties export to the rest of the world (G^{iv}). The IV estimates for β_2 equal to -0.05, -0.08, 0.09, -0.05, -0.28, -0.29, -0.14, 0.03 from column (1) to column (8), respectively. Except -0.29 from column (6) is significant at 10% confidence level, the rest of the values are insignificant.

1.4.2.2 IV Results for Instrumenting Preference Margin

Table 1.5: IV Results for Instrumenting Preference Margin

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t PREF$ $\times Pshare$	0.1013* (1.96)	0.0875* (1.66)	0.0307 (0.91)	0.1867*** (3.90)	0.4215*** (2.96)	0.2545*** (3.59)	0.1663*** (3.07)	0.0852** (2.60)
G	-0.0204 (-0.31)	-0.0383 (-0.53)	0.1469 (1.37)	0.0059 (0.04)	-0.2639 (-1.34)	-0.2783 (-1.61)	-0.0915 (-0.93)	0.0926 (0.62)
N	193,532	197,391	202,894	190,803	187,406	186,704	175,614	191,372

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

The second endogeneity concern is that the external tariffs may reversely affect the internal tariffs or both the tariffs may be impacted by some “shocks” simultaneously. Table 1.5 shows the second-stage IV result for specification (1.13), when we instrument the explanatory variable $\Delta_t PREF * Pshare$ using the instrumental variable $\Delta_t PREF^{iv} * Pshare$. This instrumental variable represents that we only instrument the tariff ($\Delta_t PREF$) of the weighted preferential tariff cut ($\Delta_t PREF * Pshare$) by the corresponding cumulative preferential tariff cut under the first FTA of the home country ($\Delta_t PREF^{iv}$). The IV estimates for β_1 equal to 0.10, 0.09, 0.03, 0.19, 0.42, 0.25, 0.17, 0.09 from column (1) to column (8). Except 0.03 from column (3) are insignificant, the rest of the columns are positive and significant. The instrumental variable for the “important” industry indicator G is the same as before. The IV estimates for β_2 equal to -0.02, -0.04, 0.15, 0.01, -0.26, -0.28, -0.09, 0.09 from column (1) to column (8). However, none of the values is significant.

1.4.2.3 IV Results for Instrumenting Import Share and Preference Margin

Table 1.6: IV Results for Instrumenting Import Share and Preference Margin

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t PREF$ $\times Pshare$	0.3237* (1.74)	0.3648** (2.54)	0.1672** (2.24)	0.3794** (2.43)	0.3789*** (2.86)	0.2736*** (3.95)	0.2775*** (3.46)	0.2405 (1.26)
G	-0.0381 (-0.62)	-0.0793 (-1.07)	0.1021 (1.03)	-0.0261 (-0.17)	-0.2579 (-1.31)	-0.2810 (-1.64)	-0.1181 (-1.20)	0.0600 (0.44)
N	193,532	197,391	202,894	190,803	187,406	186,704	175,614	191,372

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

Actually, we cannot reject the endogeneity from both the preferential tariff and the import share of FTA partners, hence, we should instrument the weighted preferential tariff cut ($\Delta_t PREF * Pshare$) using the interaction of the preferential tariff cut under the first FTA and the export share to the rest of the world. Table 1.6 presents the

second-stage IV results for the specification (1.13) when we instrument the explanatory variable $\Delta_t PREF * Pshare$ using the instrumental variable $\Delta_t PREF^{iv} * Pshare^{iv}$. The IV estimates for the coefficient β_1 of variable $\Delta_t PREF * Pshare$ equal to 0.32, 0.36, 0.17, 0.38, 0.38, 0.27, 0.28, 0.24, which are indicated in column (1) to (8), respectively. Except 0.24 from column (8) is insignificant, the other values are positive and significant. The IV estimates for the coefficient β_2 of variable G equal to -0.04, -0.08, 0.10, -0.03, -0.26, -0.28, -0.12, 0.06 from column (1) to (8), and all values are insignificant.

1.4.2.4 IV Results with Parallel FTAs

Table 1.7: IV Results with Parallel FTAs

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t PREF$	0.3237*	0.3649**	0.1559**	0.3724**	0.3826***	0.2639***	0.2691***	0.2227*
$\times Pshare$	(1.74)	(2.54)	(2.20)	(2.47)	(2.98)	(4.05)	(3.49)	(1.66)
Building	0.0188	0.0258	0.0158	0.0443	0.0466	0.0338	0.0356	0.0354
G	-0.0381	-0.0796	0.1054	-0.0325	-0.2544	-0.2763	-0.1173	0.1093
	(-0.62)	(-1.07)	(1.07)	(-0.21)	(-1.30)	(-1.62)	(-1.19)	(0.80)
Stumbling	-0.0175	-0.0336	0.0503	-0.0138	-0.1131	-0.1217	-0.0476	0.0551
N	193,532	197,391	202,894	190,803	187,406	186,704	175,614	191,372

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

Table 1.7 shows the second-stage IV results for the specification (1.14) when we instrument the explanatory variable $\Delta_t PREF * Pshare$ using $\Delta_t PREF^{iv} * Pshare^{iv}$. The only difference from the results of table 1.6 is that, in table 1.7, we control for all the subsequent FTAs that are overlapped with the current FTA period. The IV estimates for the coefficient β_1 equal to 0.32, 0.36, 0.16, 0.37, 0.38, 0.26, 0.27, 0.22, which are located in column (1) to column (8), respectively. All values are positive and significant. In general, the IV estimates for β_1 are as what we expected in section 1.3. They enhance the validity of the “building block” effect. Specifically, the coefficient 0.32 from column (1) indicates

that, given mean value of the FTA import share is 4%, a one percentage point reduction in one-year cumulative FTA tariff is associated with 0.013 (0.3237×0.04) percentage point decline in one-year cumulative MFN tariff, on average. The coefficients from other columns can be explained using the same logic. In general, when looking at the estimated coefficient β_1 from all columns, our results indicate that for an average industry and a pair of FTA partner countries, a one percentage point reduction in FTA tariff leads to a decrease in the MFN tariff ranging from 0.012 to 0.017 percentage points. The smaller elasticity in our study compared to previous literature may be attributed to the fact that we are analyzing the average import share under global FTAs, which is 0.04. If we calculate the elasticity specifically between main trade partners, it is likely to be consistent with the findings of previous papers, such as an elasticity of 0.2 from [Estevadeordal et al. \[2008\]](#) and 0.3 from [Mai and Stoyanov \[2015\]](#).

The IV estimates for the coefficient β_2 equal to -0.04, -0.08, 0.11, -0.03, -0.25, -0.28, -0.12, 0.11 from column (1) to column (8), respectively. For example, -0.04 from column (1) indicates that, on average, the 1-year cumulative MFN tariff reduction in the “important” industries is 0.04% smaller than in the counter industries. Except 0.11 from column (3) and column (8) are positive, the values from the rest columns are all negative. However, once again, none of the estimates for β_2 in columns (1) to (8) are statistically significant.

The values in the row of “Building” are calculated by multiplying the estimated coefficients β_1 of $\Delta_t P_{REF} * P_{share}$ with the corresponding mean values of cumulative preferential tariff cut ($\Delta_t P_{REF}$) and partner import share (P_{share}). The values in the row of “Stumbling” are calculated by multiplying the estimated coefficient β_2 of G with the corresponding mean values of the “important” industry indicator G . There are two purposes of generating the values from the “Building” and the “Stumbling” rows. First, it allows us to figure out that how much MFN tariff reduction is triggered by the weighted preferential tariff reduction and how much MFN reduction is smaller in the “important” industries of FTA partners in practice. Second, we are able to quantitatively compare the “building block” effect and the “stumbling block” effect. The values in the “Building” row are 0.02, 0.03, 0.02, 0.04, 0.05, 0.03, 0.04, 0.04 from column (1) to column (8). From table 1.1, we

can tell that the mean of the one-year cumulative MFN tariff reduction ($\Delta_1 MFN$) is 0.24. Therefore, over 8% of 1-year cumulative MFN tariff reduction ($\Delta_1 MFN$) is attributed to the 1-year cumulative weighted preferential tariff reduction ($\Delta_t PREF^{iv} * Pshare^{iv}$). The values in the "Stumbling" row are -0.02, -0.03, 0.05, -0.01, -0.11, -0.12, -0.05, 0.06 from column (1) to column (8). Nevertheless, none of the estimated coefficients β_2 for the variable G is significant and the estimated coefficients β_2 from column (3) and (8) are even positive. Therefore, with statistically insignificant estimates of the variable G , we cannot claim that the "stumbling block" effect exists in the worldwide trade ecosystem and the calculated values in the "Stumbling" row are proved in vain. Note that the estimated coefficients Φ_k of all the control variables $\Delta_{t-k} PREF_{cin+k} * Pshare_{cin+k}$ are not showed in table 1.7.

Table 1.8: Comparison of Different Instruments

	(1)	(2)	(3)	(4)	(5)
	OLS	$\Delta_t PREF * Pshare^{iv}$	$\Delta_t PREF^{iv} * Pshare$	$\Delta_t PREF^{iv} * Pshare^{iv}$	$\Delta_t PREF^{iv} * Pshare^{iv} + ParallelFTAs$
1-year	0.1402*	0.5046	0.1013*	0.3237*	0.3237*
2-year	0.1790*	0.3752***	0.0875*	0.3648**	0.3649**
3-year	0.0552	0.2104***	0.0307	0.1672**	0.1559**
4-year	0.2101***	0.4932*	0.1867***	0.3794**	0.3724**
5-year	0.3782**	0.5509**	0.4215***	0.3789***	0.3826***
6-year	0.2253***	0.3515***	0.2545***	0.2736***	0.2639***
7-year	0.1836***	0.3666***	0.1663***	0.2775**	0.2691***
8-year	0.1099*	0.4041	0.0852**	0.2405	0.2227*

Table 1.8 shows all the second-stage IV estimates for β_1 when we instrument the cumulative weighted preferential reduction ($\Delta_t PREF * Pshare$) using different IVs. Note that the estimated coefficients of G have poor significance levels under all three instrumenting methods, so the results for G are not included in table 1.8. The first column contains the OLS results from table 1.3, which is used here as reference.

In column (2) of table 1.8, the values are from table 1.4, where we only instrument the "share part" ($Pshare$) of the explanatory variable $\Delta_t PREF Pshare$, using the instrumental variable $\Delta_t PREF Pshare^{iv}$. It should be noted that this instrumental variable is similar to the one used in Kuenzel et al. [2019]. However, there is a difference in the approach: while

Kuenzel et al. [2019] instruments the entire weighted preferential tariff cut ($\Delta_t PREF * Pshare$) using the "export share" ($Pshare^{iv}$), we specifically use the "export share" to instrument only the corresponding "import share" ($Pshare$). The values in column (3) are from table 1.5, where we only instrument the "tariff part" ($\Delta_t PREF$) of the weighted preferential tariff reduction ($\Delta_t PREF * Pshare$), so the instrumental variable is constructed as $\Delta_t PREF^{iv} * Pshare$. The results in column (4) are exactly from table 1.6, where we use the interaction of preferential tariff reduction under the first FTA ($\Delta_t PREF^{iv}$) and the export share of FTA partners ($Pshare^{iv}$) to instrument both the "tariff part" and the "share part" of the weighted preferential tariff cut ($\Delta_t PREF * Pshare$). Therefore, the instrumental variable used here is $\Delta_t PREF^{iv} * Pshare^{iv}$.

From the comparison between column (2) (IV results) and column (1) (OLS results), it is clear that the IV results tend to be larger than the corresponding OLS results. This suggests that without properly instrumenting for the endogeneity issue related to import shares from FTA partners, the OLS results underestimate the true tariff complementarity effect. When comparing column (1) and column (3), we observe that, except for the 5-year and 6-year rows, most IV results from column (3) are slightly smaller than the corresponding OLS results from column (1). This indicates that without properly instrumenting for the endogeneity issue related to FTA tariffs, the OLS results tend to overestimate the true tariff complementarity effect.

The comparison of the IV results in column (4) with the OLS results in column (1) reveals several key findings. Firstly, the IV results consistently yield larger estimates compared to the OLS results. Secondly, the magnitudes of the IV results from column (4) generally fall within the range of the IV results from column (2) and column (3), except for the 5-year row. Thirdly, the IV results from column (4) exhibit the best significance levels among the first four columns. Lastly, the IV results from column (4) are relatively stable in magnitude. Therefore, we conclude that the results from column (4) demonstrate the effectiveness of the instrumental variable approach in addressing endogeneity and providing robust estimates. Moreover, as we are unable to reject the endogeneity concerns related to both preferential tariffs and import shares, our focus shifts to instrumenting

the weighted preferential tariff cut ($\Delta_t P_{REF} * P_{share}$) using the instrumental variables $\Delta_t P_{REF}^{iv} * P_{share}^{iv}$. After comparing the three different methods of instrumenting the explanatory variable $\Delta_t P_{REF} * share$ and identifying the most suitable approach for this project, we now turn our attention to addressing the final endogeneity issue arising from parallel FTAs.

The IV results in column (5), obtained from table 1.7, account for the influence of parallel FTAs. The only difference between the results in column (4) and column (5) is the inclusion of controls for parallel FTAs in column (5). In most cases, the results in column (5) are smaller than the corresponding results in column (4), as expected. This demonstrates that without controlling for parallel FTAs, the effect of the current FTA is overestimated. A visual comparison of the different IV results is presented in Figure 1.1. Additionally, figure 1.2 provides a visualization of the IV results from column (5) along with their corresponding 90% confidence intervals.

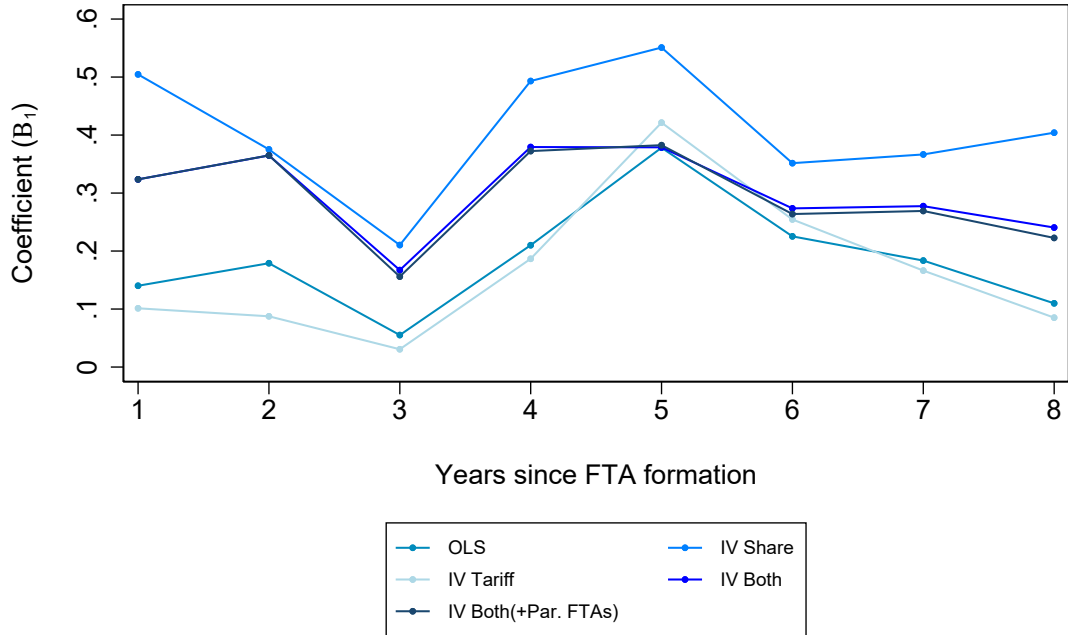


Figure 1.1: IV Results Comparison

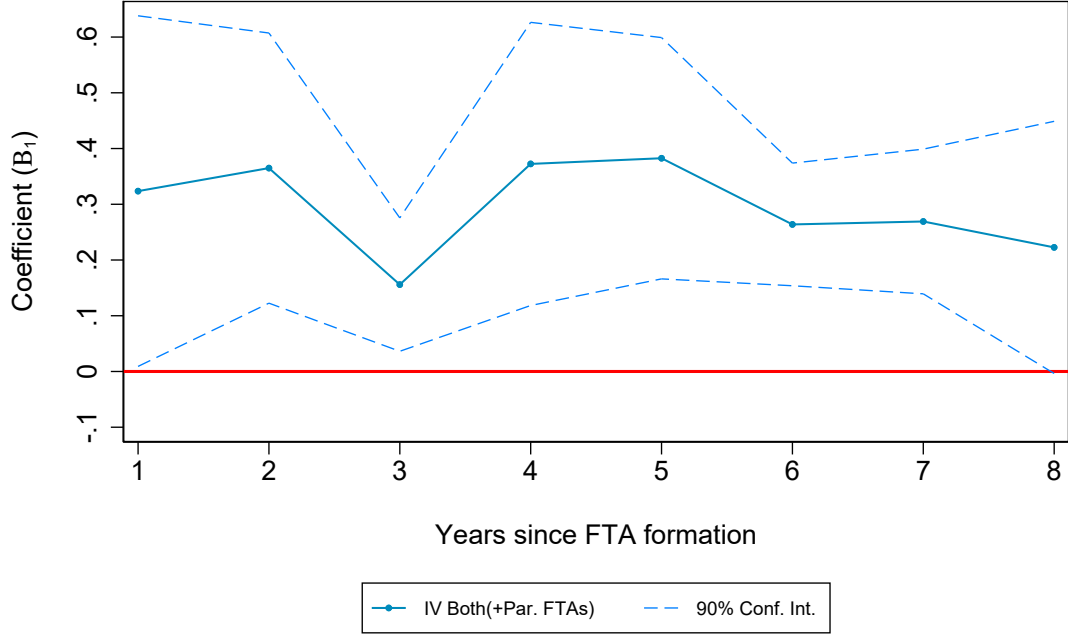


Figure 1.2: IV Result with Confidence Interval

1.5 Replications and Extensions

1.5.1 [Limão \[2006\]](#) Replication

In this section, we aim to investigate the presence of the "Stumbling block" effect, which was identified by [Limão \[2006\]](#) in the context of the US. To achieve this, we attempt to replicate the empirical results of [Limão \[2006\]](#) using their methodologies as well as the methodologies employed in this paper. It is important to note that this replication is not exact due to differences in databases and time periods used. Nonetheless, our objective is to test the underlying concept put forth by [Limão \[2006\]](#) within our own sample period and with the use of our instrumental variables.

Table 1.9 presents the second-stage IV results in the US, replicating the main specification of [Limão \[2006\]](#) by including only the industry dummy variable G . The dependent variables in columns (1) to (6) represent the 1-, 2-, 3-, 5-, 6-, and 8-year cumulative MFN reductions ($\Delta_t MFN$), with the exclusion of the 4th and 7th years due to missing preferential tariff

Table 1.9: Replication of Limão [2006]

	(1)	(2)	(3)	(4)	(5)	(6)
G	-0.0535*** (-4.86)	-0.0285** (-2.19)	-0.0230 (-1.55)	-0.1115*** (-3.15)	-0.0164 (-1.28)	-0.0526** (-1.96)
$Stumbling$	-2.1×10^{-2}	-1.1×10^{-2}	-9.2×10^{-3}	-4.1×10^{-2}	-7.8×10^{-3}	-2.4×10^{-2}
N	12,073	11,847	11,456	6,686	6,939	4,470

Notes: The dependent variables in column (1) to (6) are the cumulative MFN tariff reduction ($\Delta_t MFN$), where t equals to 1, 2, 3, 5, 6, 8, respectively. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include industry and FTA fixed effects and are clustered at industry level.

data under the first FTA formed by the US. The estimated coefficients for the dummy variable G are consistently negative, with values of -0.05, -0.03, -0.02, -0.11, -0.02, and -0.05 in columns (1) to (6). While the significance levels vary, with -0.02 being significant at the 15% confidence level in columns (3) and (5), the overall results indicate a consistently negative and significant “stumbling block” effect in the US. The next step is to examine whether this effect persists when controlling for the weighted preferential tariff cut and parallel FTAs.

Table 1.10: Replication of Limão [2006] - Our Methodology

	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta_t PREF$	0.0094	-0.1105*	0.0635	-0.1762**	0.0489	0.0083
$\times PShare$	(0.33)	(-1.83)	(0.99)	(-2.05)	(1.57)	(0.12)
$Building$	1.6×10^{-4}	-2.1×10^{-3}	1.2×10^{-3}	-2.5×10^{-3}	1.1×10^{-3}	1.7×10^{-4}
G	-0.0544*** (-4.54)	-0.0185 (-1.30)	-0.0416** (-2.47)	-0.1270*** (-3.35)	-0.0150 (-1.09)	-0.0418 (-1.39)
$Stumbling$	-2.1×10^{-2}	-7.3×10^{-3}	-1.7×10^{-2}	-4.6×10^{-2}	-7.1×10^{-3}	-1.9×10^{-2}
N	12,073	11,847	11,456	6,686	6,939	4,470

Notes: The dependent variables in column (1) to (6) are the cumulative MFN tariff reduction ($\Delta_t MFN$), where t equals to 1, 2, 3, 5, 6, 8, respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1, 2, 3, 5, 6, 8 for the corresponding column. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include industry and FTA fixed effects and are clustered at industry level.

Table 1.10 displays the second-stage IV results of specification (1.14) applied to the US case in this paper. The estimates for the interaction term $\Delta_t PREF * Pshare$ are 0.01, -0.11,

0.06, -0.17, 0.05, and 0.01 in columns (1) to (6), respectively. All the positive estimates are statistically insignificant, while -0.11 and -0.17 in columns (2) and (4) are significant. These results indicate the absence of a “building block” effect in the US. Additionally, the estimates for the industry dummy variable G are consistently negative, with values of -0.05, -0.02, -0.04, -0.13, -0.02, and -0.04 in columns (1) to (6). Although only the results in columns (1), (3), and (4) are statistically significant, the overall pattern suggests the presence of a “stumbling block” effect in the US. Hence, we cannot reject the existence of the “stumbling block” effect in the US based on these results.

1.5.2 Development Level

So far, we have confirmed that FTAs are leading to a “stumbling block” effect on MTL of the US, while resulting in a “building block” effect for the global scenario. In this section, our objective is to generalize the empirical findings and investigate whether these results are coincidental or systematic. We are particularly interested in identifying the types of countries that experience stronger “building block” or “stumbling block” effects compared to others. Given that the US is a developed country that has formed numerous FTAs, often with smaller economies, we will focus on testing three hypotheses. Firstly, based on the findings of [Estevadeordal et al. \[2008\]](#), we expect the “building block” effect to be weaker in developed countries, while the “stumbling block” effect is more likely to exist in developed countries. Secondly, we hypothesize that countries with a higher propensity to join FTAs, possibly due to their lower initial MFN tariffs, would exhibit a weaker “building block” effect. Lastly, we anticipate that if the economic size of the importer country is larger than that of the FTA exporters, we would observe a weaker “building block” effect but a stronger “stumbling block” effect in the importer country as the US case. By investigating these hypotheses, our aim is to uncover systematic patterns regarding the effects of FTAs on different countries and provide further insights into the dynamics of trade relationships.

Therefore, we extend the regression model to examine whether the effect of FTAs on MTL varies across different categories of countries. We classify all countries in the data sample into three categories based on different dimensions. To incorporate these categories

into the analysis, we modify specification (1.14) by introducing additional interaction terms. By doing so, we aim to investigate the heterogeneous effects of FTAs on MTL across different country groups and shed light on the factors that contribute to these variations.

$$\begin{aligned}
\Delta_t MFN_{cin} = & \beta_0 + \beta_1 \Delta_t PREF_{cin} * Pshare_{cin} + \beta_2 d_2 * \Delta_t PREF_{cin} * Pshare_{cin} + \\
& \beta_3 d_3 * \Delta_t PREF_{cin} * Pshare_{cin} + \beta_4 G_{cin} + \beta_5 d_2 * G_{cin} + \beta_6 d_3 * G_{cin} + \\
& \sum_{k=1}^{t-1} \Phi_k \Delta_{t-k} PREF_{cin+k} * Pshare_{cin+k} + \alpha_c + \alpha_i + \alpha_n + \varepsilon_{cin} \quad (1.15)
\end{aligned}$$

Our first hypothesis is that FTAs play different roles in countries with different development levels. To test this hypothesis, we categorize the countries in our sample based on their development levels and construct three dummy variables: d_1 (omitted), d_2 , and d_3 . The dummy variable d_2 takes the value of 1 if the importer is a least developed country, while d_3 takes the value of 1 if the importer is a developed country. The development level of a country is determined by its national income level in the year before the formation of each FTA. Based on this categorization, we expect to observe the following patterns in the regression results: a positive and significant coefficient (β_1) for the interaction term between the cumulative preferential tariff reduction and the import share of FTA partners in least developed countries (d_2), an insignificant coefficient (β_2) for the interaction term involving the "important" industry indicator (G), a negative and significant coefficient (β_3) for the interaction term in developed countries (d_3), and insignificant coefficients (β_4 and β_5) for the other interaction terms. Additionally, we expect a negative and significant coefficient (β_6) for the "important" industry indicator (G) in the overall model. By analyzing these coefficients, we can assess the differential effects of FTAs on MTL across countries with different development levels.

Table 1.11 contains the second-stage IV results for the specification (1.15) in this case. The estimates for the coefficients (β_1) of the variable $\Delta PREF_{cin} * Pshare_{cin}$ from column (1) to column (8) are 0.39, 0.27, 0.13, 0.41, 0.85, 0.97, 0.23, 0.02, respectively. All the coefficients are positive and 0.27, 0.97 and 0.23 from column (2), (6) and (7) are significant.

Table 1.11: Development

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t PREF$	0.3883	0.2668**	0.1275	0.4103	0.8523	0.9749**	0.2275**	0.0185
$\times PShare$	(0.75)	(2.58)	(1.56)	(1.15)	(1.53)	(2.41)	(2.37)	(0.13)
$d_2 \Delta_t PREF$	-0.2215	0.3723	-0.0692	-0.0799	-1.3372*	-0.7145*	0.1355	0.0939
$\times PShare$	(-0.32)	(1.02)	(-0.25)	(-0.21)	(-1.73)	(-1.74)	(1.63)	(0.61)
$Sum(\beta_1 + \beta_2)$	[0.29]	[1.65]	[0.23]	[2.35]	[-0.96]	[2.54]	[6.03]	[1.75]
$d_3 \Delta_t PREF$	-0.0704	0.5729	1.0940**	2.2738***	-0.5110	-0.7835*	-0.0121	0.3697
$\times PShare$	(-0.13)	(1.21)	(2.17)	(2.74)	(-0.90)	(-1.88)	(-0.08)	(1.04)
$Sum(\beta_1 + \beta_3)$	[1.59]	[1.82]	[2.45]	[3.58]	[2.65]	[3.35]	[1.68]	[1.08]
G	-0.1735	-0.3176*	-0.0040	0.0186	-0.6123**	-0.4708	-0.2541	0.3130
	(-1.08)	(-1.91)	(-0.02)	(0.05)	(-2.12)	(-1.58)	(-0.68)	(0.54)
$d_2 G$	0.1928	0.0483	-0.0447	-0.1462	0.5725	0.0500	0.0412	-0.3717
	(0.97)	(0.28)	(-0.15)	(-0.27)	(1.60)	(0.14)	(0.10)	(-0.61)
$Sum(\beta_4 + \beta_5)$	[0.10]	[-1.39]	[-0.23]	[-0.29]	[-0.15]	[-1.83]	[-0.93]	[-0.32]
$d_3 G$	0.1571	0.3026*	0.1500	-0.0919	0.4423	0.2449	0.1908	-0.1976
	(1.13)	(1.70)	(0.57)	(-0.27)	(1.58)	(0.87)	(0.54)	(-0.33)
$Sum(\beta_4 + \beta_6)$	[-0.26]	[-0.21]	[1.26]	[-0.48]	[-0.90]	[-1.30]	[-1.04]	[0.79]
N	193,532	197,391	202,894	190,803	187,406	186,704	175,614	191,372

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. d_2, d_3 are the dummy variables indicating that if the country is a developing or developed country. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. t-values for summation coefficient in bracket. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

For example, 0.27 from column (2) implies that, if we fixed the import share of FTA partners at its mean value, on average, every 1 percent 2-year cumulative reduction in the weighted preferential tariff leads to a 0.27 percent 2-year cumulative reduction in the MFN tariff among the least developed countries. In general, since all significant estimates are positive, we conclude that the “building block” effect exists in the least developed countries. The estimates for the coefficient β_2 of the variable $d_2 * \Delta PREF_{cin} * Pshare_{cin}$ are -0.22, 0.37, -0.07, -0.08, -1.34, -0.71, 0.14, 0.09 from column (1) to (8), respectively. Only -1.34 and -0.71 from column (5) and (6) are significant. Technically, -0.71 from column (6) indicates that, on average, if we fixed the import share of FTA partners at its mean value, every 1 percent 6-year cumulative reduction in preferential tariff leads to a 0.71 percent 6-year cumulative reduction in the MFN tariff deeper in the least developed countries than the developing

countries. The estimates for the coefficient β_3 of the variable $d_3 * \Delta PREF_{cin} * Pshare_{cin}$ are -0.07, 0.57, 1.09, 2.27, -0.51, -0.78, -0.01, 0.37 from column (1) to (8), respectively. Except 1.09, 2.27 and -0.78 are from column (3), (4) and column (6), the other columns are all insignificant. The -0.78 from column (6) shows that, on average, every 1 percent of the 6-year cumulative reduction in preferential tariff leads to a 0.78 percent 6-year cumulative reduction in the MFN tariff deeper in the least developed countries than the developed countries if we fixed the import share of FTA partners. Since the significant estimated coefficients of β_2 and β_3 are changing in signs from column (1) to (8), we cannot observe the clear pattern of the strength of the “building block” effect among different development levels. The values in the “Sum($\beta_1 + \beta_2$)” and “Sum($\beta_1 + \beta_3$)” rows are the t-values for the “coefficient” for the developing and the developed countries. Since most values from these two rows are positive and significant, we conclude that the “building block” effect also exists in the developing and the developed countries. At the bottom part of the table, the estimates for the variables G , $d_2 * G$ and $d_3 * G$ are varying in signs and have the poor significance levels. The t-values for the summation coefficients $\beta_4 + \beta_5$ and $\beta_4 + \beta_6$ are mostly insignificant. Therefore, we cannot find out the “stumbling block” effect in the least developed, developing or the developed countries.

In summary, while we observe the “building block” effect in least developed countries, the results do not show a clear pattern of the effect’s strength across different development levels. The “building block” effect seems to exist in all categories of countries, regardless of their development status. On the other hand, the “stumbling block” effect is not evident in any of the country categories examined.

1.5.3 Propensity of Joining FTAs

The second hypothesis is that the “building block” effect is weaker in the countries having higher propensity towards joining FTAs. The logic of this hypothesis is complicated. Countries that demand in signing more FTAs are prone to keep their borders open, so there should be a strong tariff complementarity effect. However, as more FTAs have been signed, there would be less room for further tariff reductions. In order to test this hypothesis, we

need to change the definition of d from the development case. We first count the total number of effective FTAs that each country has formed by the end of year 2017, which is the last year of the study period. Secondly, we divide the sample by the median value of the distribution of the total number of FTAs, so the two groups represent the countries with larger number of FTAs and smaller number of FTAs. Lastly, the dummy variable d takes the value of one if a country belongs to the group of countries with the larger number of FTAs or zero otherwise. We consider that the larger number of FTAs formed by a country, the less room for tariff variation in this country. Therefore, we expect to obtain the positive and significant β_1 , negative β_2 , insignificant β_3 , negative and significant β_4 .

Table 1.12: FTA Propensity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t Pmargin$	0.1562*	0.2435**	0.4258*	0.1534	1.2017***	0.4367*	0.2480*	1.2296***
$\times ImShare_t$	(1.97)	(2.63)	(1.81)	(1.41)	(6.35)	(1.71)	(1.82)	(4.63)
$d \times \Delta_t Pmargin$	1.2649***	0.6299	-0.0084	-0.0100	-1.1103***	-0.3199	-0.3012	-1.1852***
$\times ImShare_t$	(5.18)	(1.12)	(-0.02)	(-0.08)	(-5.19)	(-1.23)	(-1.21)	(-3.49)
<i>Summation</i>	1.4211***	0.8734	0.4174	0.1434**	0.0914	0.1168*	-0.0533	0.0445
	(6.05)	(1.53)	(1.23)	(2.40)	(1.58)	(1.87)	(-0.25)	(0.29)
G	-0.0399	-0.0038	0.7079	0.2584	-0.9434	-0.0238	-0.0608	-0.4515
	(-0.61)	(-0.05)	(1.29)	(1.62)	(-1.54)	(-0.14)	(-0.29)	(-1.05)
$d \times G$	0.0199	0.0963	-0.4633	-0.3979**	0.6772	-0.0831	0.0267	0.3446
	(0.32)	(0.86)	(-1.10)	(-2.10)	(1.15)	(-0.44)	(0.14)	(-0.1069)
<i>Summation</i>	-0.0200	0.0926	0.2446	-0.1394	-0.2662	-0.1069	-0.0341	-0.1069
	(-0.26)	(0.94)	(1.32)	(-0.77)	(-1.06)	(-0.70)	(-0.23)	(-0.42)
N	193,532	222,799	238,152	281,970	285,956	299,232	251,539	299,743

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. d is the dummy variable indicating that if the country has higher propensity of joining FTAs. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

In table 1.12, the estimates of coefficient β_1 and β_2 demonstrate that the “building block” effect exists in the countries with relatively smaller number of FTAs and the effect is weaker in the other group of countries. The values in the “Summation” ($\beta_1 + \beta_2$) row indicate that the “building block” effect is absent in the countries with relatively larger number of FTAs. This is consistent with the replication results of the US case, since the

US belongs to the group of countries with the comparatively large number of FTAs. Like results indicated from table 1.12, the estimates for the coefficients β_3 , β_4 and the summation $(\beta_3 + \beta_4)$ are also varying in signs and have poor significance levels. Therefore, the strength of the “stumbling block” effect is not directly correlated to a country’s propensity towards forming FTAs.

1.5.4 Relative Economic Size

Last but not least, in order to test the third hypothesis, we classify countries by their GDP level relative to that of exporters. First, calculate the relative GDP ratio (GDP of importer relative to the total GDP of exporters under the same FTA). Second, split the sample by the median value of the distribution of the GDP ratio. Last, construct the dummy variable d taking the value of one if a country has larger relative GDP ratio or zero otherwise. So intuitively, d equals to one representing that the importer has relatively larger economic size as compare to the exporters under the same FTA. Along with the explanation from Limão [2006], we expect to obtain the positive and significant β_1 , negative β_2 , insignificant β_3 , negative and significant β_4 .

In table 1.13, if we focus on the last 4 columns, the estimates of the coefficients β_1 and β_2 show that the FTAs have weaker “building block” effect if they are formed by an importer with the relatively larger economic size than that of the FTA exporters. The insignificant values from “Summation” $(\beta_1 + \beta_2)$ demonstrate that the “building block” effect is absent in the countries which have larger economic sizes than those of the FTA exporters. However, due to the poor significance levels of the coefficients β_3 , β_4 and summation $(\beta_3 + \beta_4)$, the evidence for the “stumbling block” effect is still not explicit.

1.6 Conclusion

In the field of trade liberalization, the question of whether forming a Free Trade Agreement (FTA) accelerates or decelerates multilateral trade liberalization is still subject to debate. In this paper, we aim to contribute to this discussion by analyzing the overall effects of FTAs

Table 1.13: Relative Economic Size

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta_t Pmargin$	0.1641**	0.3002***	0.1682*	0.1561*	1.0491***	0.3930*	0.2315*	1.2119***
$\times ImShare_t$	(2.12)	(2.79)	(1.72)	(1.92)	(4.25)	(1.91)	(1.93)	(4.66)
$d \times \Delta_t Pmargin$	0.4614	0.0795	1.1685***	0.0927	-0.4312	-0.2316	-0.1801	-1.3261***
$\times ImShare_t$	(1.25)	(0.30)	(4.59)	(0.50)	(-0.47)	(-0.90)	(-0.54)	(-4.66)
<i>Summation</i>	0.6255*	0.3798	1.3367***	0.2488	0.6178	0.1614	0.0514	-0.1142
	(1.75)	(1.48)	(5.00)	(1.55)	(0.71)	(0.97)	(0.16)	(-1.18)
G	-0.0656	0.0127	0.3359	0.0601	-0.7936	-0.1417	-0.1201	-0.2335
	(-0.85)	(0.16)	(1.08)	(0.39)	(-1.51)	(-0.94)	(-0.79)	(-0.68)
$d \times G$	0.0603	0.0958	0.2649	-0.0585	0.4702	0.0936	0.1670	-0.0084
	(0.74)	(0.82)	(1.09)	(-0.19)	(0.99)	(0.51)	(0.92)	(-0.02)
<i>Summation</i>	-0.0052	0.1085	0.6008	0.0016	-0.3234	-0.0481	0.0499	-0.2419
	(-0.07)	(1.08)	(1.34)	(0.01)	(-0.91)	(-0.29)	(0.26)	(-0.97)
N	193,532	222,799	238,152	281,970	285,956	299,232	251,539	299,743

Notes: The dependent variables in column (1) to (8) are the 1- to 8-year cumulative MFN tariff reduction ($\Delta_t MFN$), respectively. $\Delta_t PREF \times Pshare$ is varying as t equals 1 to 8 for the corresponding column. d is the dummy variable indicating that if the FTA is formed by a importer has higher GDP than exporters. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country, industry and FTA fixed effects and are double clustered at country and industry level.

on global Multilateral Trade Liberalization. Our analysis combines two commonly studied channels from previous literature: the “Political Economy” effect proposed by [Limão \[2006\]](#) and the “Complementarity” effect described by [Ornelas \[2005\]](#).

Our findings provide evidence for a relatively consistent causal relationship between the reduction in weighted preferential tariffs and MFN (Most Favored Nation) tariffs, suggesting the presence of a “building block” effect of FTAs on global trade liberalization. Specifically, we find that, on average, a one percentage point reduction in FTA tariffs is associated with a decrease of 0.012 to 0.017 percentage points in MFN tariffs for an average industry and a pair of FTA partner countries. It is worth noting that our estimated elasticity values are relatively smaller compared to previous literature. This discrepancy can be attributed to our consideration of the average import share under global FTAs, which is approximately 0.04. However, when examining the elasticity specifically between main trade partners, our results align with the findings of previous literature.

However, we did not find sufficient evidence to support the existence of a “stumbling

block” effect. Furthermore, we observe that the strength of the “building block” effect varies in certain contexts. For instance, it is relatively weaker when the FTA is formed by developed countries, countries with a higher number of FTAs, or importers with larger economic size compared to exporters.

These findings indirectly shed light on the “stumbling block” effect observed in the US, as the country is developed and has formed numerous FTAs, often with smaller economies. As the number of FTAs continues to increase worldwide, it becomes crucial to examine the integrated effects of FTAs on global trade liberalization. By providing empirical evidence on these effects, our study contributes to a better understanding of the complex dynamics between FTAs and multilateral trade liberalization.

References

- O. Al-Ubaydli and P. A. McLaughlin. RegData: A numerical database on industry-specific regulations for all United States industries and federal regulations, 1997–2012. *Regulation & Governance*, 11(1):109–123, March 2017. doi: 10.1111/rego.12107. URL <https://ideas.repec.org/a/wly/reggov/v11y2017i1p109-123.html>.
- K. Bagwell and R. W. Staiger. Regionalism and multilateral tariff co-operation. In *International trade policy and the pacific rim*, pages 157–190. Springer, 1999.
- J. B. Bailey and D. W. Thomas. Regulating away competition: the effect of regulation on entrepreneurship and employment. *Journal of Regulatory Economics*, 52(3):237–254, December 2017. doi: 10.1007/s11149-017-9343-9.
- J. B. Bailey, D. W. Thomas, and J. R. Anderson. Regressive effects of regulation on wages. *Public Choice*, 180(1/2):pp. 91–103, 2019. ISSN 00485829, 15737101. URL <https://www.jstor.org/stable/48704155>.
- D. Chambers and C. O’Reilly. Regulation and income inequality in the United States. *European Journal of Political Economy*, 72(C), 2022. doi: 10.1016/j.ejpoleco.2021.1. URL <https://ideas.repec.org/a/eee/poleco/v72y2022ics0176268021000872.html>.
- D. Chambers, C. A. Collins, and A. Krause. How do federal regulations affect consumer prices? an analysis of the regressive effects of regulation. *Public Choice*, 180:57–90, 2019.
- P.-L. Chang and M.-J. Lee. The wto trade effect. *Journal of International Economics*, 85(1):53–71, 2011.

- B. Coffey, P. A. McLaughlin, and P. Peretto. The cumulative cost of regulations. *Review of Economic Dynamics*, 38:1–21, 2020. ISSN 1094-2025. doi: <https://doi.org/10.1016/j.red.2020.03.004>. URL <https://www.sciencedirect.com/science/article/pii/S1094202520300223>.
- T. S. Eicher and C. Henn. In search of wto trade effects: Preferential trade agreements promote trade strongly, but unevenly. *Journal of International Economics*, 83(2):137–153, 2011.
- A. Estevadeordal, C. Freund, and E. Ornelas. Does regionalism affect trade liberalization toward nonmembers? *The Quarterly Journal of Economics*, 123(4):1531–1575, 2008.
- S. Hakobyan. Accounting for underutilization of trade preference programs: The us generalized system of preferences. *Canadian Journal of Economics/Revue canadienne d'économique*, 48(2):408–436, 2015.
- B. Herz and M. Wagner. The dark side of the generalized system of preferences. *Review of International Economics*, 19(4):763–775, 2011.
- B. Karacavaali and N. Limão. The clash of liberalizations: preferential vs. multilateral trade liberalization in the european union. *Journal of international Economics*, 74(2):299–327, 2008.
- T. D. Ketterer, D. Bernhofen, and C. Milner. Preferences, rent destruction and multilateral liberalization: The building block effect of cusfta. *Journal of International Economics*, 92(1):63–77, 2014.
- T. D. Ketterer, D. M. Bernhofen, and C. Milner. The impact of trade preferences on multilateral tariff cuts: Evidence for japan. *Journal of the Japanese and International Economies*, 38:31–51, 2015.
- D. J. Kuenzel, R. R. Sharma, et al. *Preferential Trade Agreements and MFN Tariffs: Global Evidence*. Wesleyan University, Department of Economics, Public Affairs Center, 2019.

- N. Limão. Preferential trade agreements as stumbling blocks for multilateral trade liberalization: Evidence for the united states. *American Economic Review*, 96(3):896–914, 2006.
- D. S. Lucas and C. J. Boudreaux. National regulation, state-level policy, and local job creation in the united states: A multilevel perspective. *Research Policy*, 49(4):103952, 2020. ISSN 0048-7333. doi: <https://doi.org/10.1016/j.respol.2020.103952>. URL <https://www.sciencedirect.com/science/article/pii/S0048733320300329>.
- J. Mai and A. Stoyanov. The effect of the canada-us free trade agreement on canadian multilateral trade liberalization. *Canadian Journal of Economics/Revue canadienne d'économique*, 48(3):1067–1098, 2015.
- M. S. Manger and K. C. Shadlen. Political trade dependence and north–south trade agreements. *International studies quarterly*, 58(1):79–91, 2014.
- E. Ornelas. Rent destruction and the political viability of free trade agreements. *The Quarterly Journal of Economics*, 120(4):1475–1506, 2005.
- E. Ornelas and M. Ritel. The not-so-generalised effects of the generalized system of preferences. *The World Economy*, 43(7):1809–1840, 2020.
- Ç. Özden and E. Reinhardt. The perversity of preferences: Gsp and developing country trade policies, 1976–2000. *Journal of Development Economics*, 78(1):1–21, 2005.
- A. Postigo. Explaining utilization of free trade agreements by sectoral interests and binding of unilateral concessions. *International Development*, 148:1–63, 2013.
- A. K. Rose. Do we really know that the wto increases trade? *American Economic Review*, 94(1):98–114, March 2004. doi: 10.1257/000282804322970724. URL <https://www.aeaweb.org/articles?id=10.1257/000282804322970724>.
- A. Subramanian and S.-J. Wei. The wto promotes trade, strongly but unevenly. *Journal of international Economics*, 72(1):151–175, 2007.

M. Tomz, J. L. Goldstein, and D. Rivers. Do we really know that the wto increases trade?
comment. *American Economic Review*, 97(5):2005–2018, 2007.

Figures



Figure 1.3: Number of Effective FTAs (1994-2017)

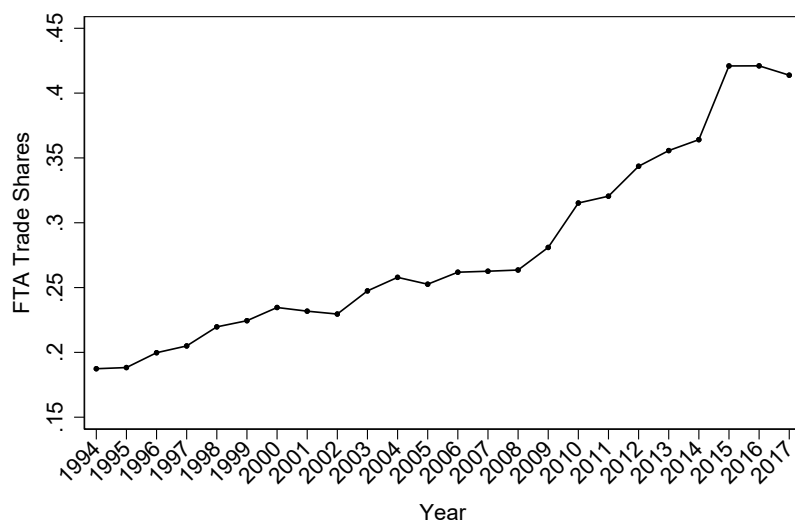


Figure 1.4: Trade under FTA over Global Trade (1994-2017)

Appendix A: Chapter 1

In this section, we will briefly prove all the equilibrium equations which appear in the theory section. To have a clearer picture, we further simplify the notation and assume that besides the numeraire, there is only one type of good trading between the 3 countries. Hence, the preference parameter $d = 1$. Therefore, we have

Aggregate utility:

$$U = X_0 + \ln X;$$

Aggregate consumption:

$$X = \left(x_H^{\frac{\sigma-1}{\sigma}} + x_P^{\frac{\sigma-1}{\sigma}} + x_N^{\frac{\sigma-1}{\sigma}} \right)^{\frac{\sigma}{\sigma-1}};$$

Aggregate price index:

$$P = (p_H^{1-\sigma} + p_P^{1-\sigma} + p_N^{1-\sigma})^{\frac{1}{1-\sigma}};$$

Individual demand function maximizing the utility:

$$x_c = p_c^{-\sigma} (p_H^{1-\sigma} + p_P^{1-\sigma} + p_N^{1-\sigma})^{-1}$$

$$x_c = I p_c^{-\sigma} P^{\sigma-1} \quad \forall c \in \{H, P, N\},$$

where p_c is the individual price.

We take the constant elasticity of substitution aggregate form on both sides of the individual demand function above, it becomes

$$X = I P^{-\sigma} P^{\sigma-1} = \frac{I}{P},$$

since the individual level of demand and price turn to the aggregate level. Note that I has normalized to 1 in the theory section, and $I = PX$ is rational in this section since all

income has spent on the only good at the equilibrium level.

Profit function:

$$\pi_c = \left(\frac{p_c}{1 + \tau_c} - c_c \right) q_c \quad \forall c \in \{H, P, N\},$$

we use q_c here since the profit function has not yet reached the equilibrium level. $\tau_H = 0$ since the tariff is imposed by the home country.

Profit-maximizing price strategy:

$$p_c = \left(\frac{\sigma}{\sigma - 1} \right) c_c (1 + \tau_c) \quad \forall c \in \{H, P, N\};$$

Equilibrium profit:

$$\pi_c = \frac{p_c x_c}{\sigma (1 + \tau_c)} \quad \forall c \in \{H, P, N\};$$

Aggregate government objective function:

$$G(\tau) = CS(\tau) + PS(\tau) + TR(\tau),$$

where τ is MFN tariff.

Consumer surplus:

$$CS(\tau) = U(X_0, X) - \sum_{c \in \{H, P, N\}} p_c x_c = X_0 + \ln X - I;$$

Producer surplus:

$$PS(\tau) = \pi_H = \frac{p_H x_H}{\sigma};$$

Tariff Revenue:

$$TR(\tau) = \frac{\tau}{1 + \tau} p_N x_N + \frac{\tau_P}{1 + \tau_P} p_P x_P.$$

Now we calculate the derivatives and obtain the following results:

Derivative of consumer surplus:

$$\frac{\partial CS(\tau)}{\partial \tau} = \frac{\partial \ln X}{\partial \tau} = -\frac{1}{P} \frac{\partial P}{\partial \tau} = -X \frac{\partial P}{\partial \tau};$$

Derivative of producer surplus:

$$\frac{\partial PS(\tau)}{\partial \tau} = \frac{p_H}{\sigma} \frac{\partial x_H}{\partial \tau} = \left(\frac{\sigma - 1}{\sigma} \right) p_H p_H^{-\sigma} P^{-1} \frac{\partial P}{\partial \tau} = \left(\frac{\sigma - 1}{\sigma} \right) p_H p_H^{-\sigma} \frac{X}{I} \frac{\partial P}{\partial \tau} = \left(\frac{\sigma - 1}{\sigma} \right) s_H X \frac{\partial P}{\partial \tau},$$

where $s_H = \frac{p_H x_H}{I}$ is the domestic market share supplied by domestic good.

For better understanding, we need calculate some intermediate results at equilibrium level before deriving the tariff revenue. First we have the derivative of the price index with respect to MFN tariff,

$$\frac{\partial P(\tau)}{\partial \tau} = P^\sigma p_c^{-\sigma} \frac{\partial p_c}{\partial \tau};$$

therefore,

$$\frac{\partial p_c}{\partial \tau} = \frac{\partial P(\tau)}{\partial \tau} \left(\frac{p_c}{P} \right)^\sigma.$$

Also, by substituting the equilibrium marginal cost, we get

$$\frac{\partial p_c}{\partial \tau} = c_c \left(\frac{\sigma}{\sigma - 1} \right) = \frac{p_c}{1 + \tau} \quad \forall c \in \{P, N\}.$$

Recall that $PX = I = 1$, hence

$$p_c x_c = \frac{p_c x_c}{PX} PX \frac{\partial P}{\partial \tau} \left(\frac{p_c}{P} \right)^\sigma \left(\frac{\partial p_c}{\partial \tau} \right)^{-1} = s_c \frac{1}{s_c} (1 + \tau) X \frac{\partial P}{\partial \tau} = (1 + \tau) X \frac{\partial P}{\partial \tau},$$

where s_c is defined similarly as s_H .

Thus, the derivative of tariff revenue

$$\frac{\partial TR(\tau)}{\partial \tau} = \frac{1}{(1 + \tau)^2} p_N x_N + \frac{\tau}{1 + \tau} \left(x_N \frac{\partial p_N}{\partial \tau} + p_N \frac{\partial x_N}{\partial \tau} \right) + \frac{\tau_P}{1 + \tau_P} p_P \frac{\partial x_P}{\partial \tau},$$

where the first set of terms on the right-hand side of the equation is

$$\frac{1}{(1+\tau)^2} p_N x_N = \frac{1}{(1+\tau)} X \frac{\partial P}{\partial \tau};$$

we have,

$$x_N \frac{\partial p_N}{\partial \tau} = x_N \frac{\partial P}{\partial \tau} \left(\frac{p_N}{P} \right)^\sigma = \frac{p_N x_N}{P X} \left(\frac{p_N}{P} \right)^{\sigma-1} X \frac{\partial P}{\partial \tau} = s_N \frac{1}{s_N} X \frac{\partial P}{\partial \tau} = X \frac{\partial P}{\partial \tau},$$

$$p_N \frac{\partial x_N}{\partial \tau} = -\sigma x_N \frac{\partial p_N}{\partial \tau} + (\sigma - 1) \frac{p_N x_N}{P} \frac{\partial P}{\partial \tau} = -\sigma X \frac{\partial p_N}{\partial \tau} + (\sigma - 1) s_N X \frac{\partial P}{\partial \tau},$$

then the second set of terms on the right-hand side of the equation becomes

$$\frac{\tau}{1+\tau} \left(x_N \frac{\partial p_N}{\partial \tau} + p_N \frac{\partial x_N}{\partial \tau} \right) = \frac{\tau}{1+\tau} [(\sigma - 1) s_N - (\sigma - 1)] X \frac{\partial P}{\partial \tau};$$

the third set of terms on the right-hand side of the equation is

$$\frac{\tau_P}{1+\tau_P} p_P \frac{\partial x_P}{\partial \tau} = \frac{\tau_P (\sigma - 1)}{1+\tau_P} \frac{p_P x_P}{P} \frac{\partial P}{\partial \tau} = \frac{\tau_P (\sigma - 1)}{1+\tau_P} s_P X \frac{\partial P}{\partial \tau},$$

therefore,

$$\frac{\partial TR(\tau)}{\partial \tau} = X \frac{\partial P}{\partial \tau} \left\{ \frac{1}{(1+\tau)} + \frac{\tau}{1+\tau} [(\sigma - 1) s_N - (\sigma - 1)] + \frac{\tau_P (\sigma - 1)}{1+\tau_P} s_P \right\}.$$

Derivative of tariff revenue in absence of FTA:

$$\frac{\partial TR(\tau)}{\partial \tau} = X \frac{\partial P}{\partial \tau} \left\{ \frac{1}{(1+\tau)} + \frac{\tau}{1+\tau} [(\sigma - 1) (s_N + s_P) - (\sigma - 1)] \right\};$$

Derivative of tariff revenue in presence of FTA:

$$\frac{\partial TR(\tau)}{\partial \tau} = X \frac{\partial P}{\partial \tau} \left\{ \frac{1}{(1+\tau)} + \frac{\tau}{1+\tau} [(\sigma-1)s_N - (\sigma-1)] + \frac{\tau_P(\sigma-1)}{1+\tau_P} s_P \right\};$$

Import demand elasticity:

$$\varepsilon = \frac{\partial x_c}{\partial p_c} \frac{p_c}{x_c} = \left[-\sigma \frac{q_c}{p_c} + (\sigma-1) \frac{q_c}{p_c} \frac{\partial P}{\partial p_c} \right] \frac{p_c}{x_c} = -\sigma + (\sigma-1) \left(\frac{p_c}{P} \right)^{1-\sigma} = -\sigma + (\sigma-1) s_c,$$

where s_c is different in the two cases. $s_c = (s_P + s_N)$ when there is no FTA and $s_c = s_N$ when there is an FTA.

Derivative of government objective function:

$$\frac{\partial TR(\tau)}{\partial \tau} = X \frac{\partial P}{\partial \tau} \left\{ -1 + \left(\frac{\sigma-1}{\sigma} \right) s_H + (\sigma-1) \left[\frac{\tau}{1+\tau} (s_N - 1) + \frac{\tau_P}{1+\tau_P} s_P \right] + \frac{1}{1+\tau} \right\} = 0;$$

The optimal MFN tariff without FTA becomes

$$\varepsilon \frac{\tau}{1+\tau} = \frac{\sigma-1}{\sigma} s_H;$$

The optimal MFN tariff with FTA becomes

$$\varepsilon \frac{\tau}{1+\tau} = \frac{\sigma-1}{\sigma} s_H + (\sigma-1) s_P \frac{\tau_P}{1+\tau_P}.$$

Chapter 2

Did the intended beneficiaries of the Generalized System of Preference (GSP) benefit from it?

2.1 Introduction

The Generalized System of Preference (GSP), endorsed by the World Trade Organization (WTO), is a trade program that encourages developed countries to grant preferential tariff treatment to developing countries without expecting reciprocal concessions in return. The primary objective of the GSP is to promote economic growth, foster export diversification, and alleviate poverty in eligible developing countries by granting them improved market access to developed countries. It is widely acknowledged that developing countries often have limitations in terms of production capacities, resources, and competitiveness compared to their developed counterparts in the world trading system. Therefore, the GSP program is seen as a justified initiative that supports the growth of emerging industries in developing countries, even if it involves temporary compromises in terms of the economic efficiency and non-discrimination principle in the international trading system. Interestingly, there is ongoing debate and differing opinions within the literature regarding the effectiveness of

the GSP program.

Our aim is to contribute to the ongoing debate by investigating the underlying motives of donor countries in determining their GSP preferences. Using observed variation in the GSP tariff rates, we examine the degree to which the donor countries internalize the impact of the program on welfare of recipient countries. Our findings reveal limited evidence supporting the notion that the GSP preferences are set to maximize welfare improvement of the developing countries. This suggests that the GSP program falls short of utilizing fully its potential, as the concerns for the welfare of recipient countries by donor countries are overshadowed by other trade policy considerations.

The GSP program has a rich history that dates back to 1971 when it was established as part of the General Agreement on Tariffs and Trade (GATT). Originally intended as a temporary measure, it provided GATT contracting parties with the ability to grant preferential tariff rates to developing countries for a period of ten years. However, in 1979, following the Tokyo Round of multilateral trade negotiations, the temporary tariff treatment became permanent. Under the GSP program, developed countries have the legal authority to deviate from the Most Favored Nation (MFN) principle and apply more favorable tariffs to developing countries, without extending the same benefits to other developed nations. As a result, most industrialized nations have introduced their own GSP schemes, extending preferential treatment to eligible developing countries as well as least developed countries. Furthermore, specific emerging economies have implemented their own GSP programs targeting least developed countries (LDCs). Despite the widespread adoption of the program, its efficiency has faced significant criticism and skepticism. Various concerns have been raised, including the substantial discretion exercised by GSP donor countries in the selection of GSP-eligible beneficiaries and industries, the relatively high GSP tariffs imposed on sectors such as agriculture and textiles, and the stringent rules of origin requirements that hinder the participation of developing countries in global value chains.

The existing body of literature has extensively investigated the effectiveness of the GSP from various perspectives. Several studies have examined the impact of GSP status on the

export volume of eligible countries (Rose [2004]; Subramanian and Wei [2007]; Eicher and Henn [2011]). Others analyzed the influence of GSP status on the external trade policies of GSP donors or recipients (Özden and Reinhardt [2005]; Ketterer et al. [2015]). Studies by Postigo [2013] and Hakobyan [2015] have assessed the utilization rate and uncertainty associated with GSP tariff rates. However, divergent conclusions have emerged across these studies, primarily due to variations in the measurement of GSP efficiency. For instance, Hakobyan [2015] discovered that 40% of American GSP-eligible imports did not claim the GSP preference. The author attributed the lower-than-expected utilization rate of preference to factors such as the beneficiaries’ production structure and access to alternative preferential programs.

Given the lack of consensus on the efficiency of the GSP, it is evident that further investigation is necessary to obtain a more comprehensive understanding of whether the implementation of the GSP program genuinely aims to benefit its intended recipients. Therefore, it is crucial to delve into the underlying determinants of GSP efficiency, specifically examining whether GSP donors establish their GSP rates in a manner that maximizes trade opportunities for recipients, aligning with the program’s intended objectives. Surprisingly, despite its significance, this particular aspect has received relatively limited attention within the existing trade literature.

In this project, our objective is to contribute to the existing research by examining the factors affecting country’s choice of GSP. We develop a simple theoretical model of endogenous GDP tariff choice that allows us to analyze the determinants of tariff rates. A key aspect of our model is the inclusion of the assumption that donor governments not only aim to maximize their national welfare but also take into consideration the impact of GSP tariffs on the profits of firms from recipient countries. This assumption reflects the idea that donor governments internalize the effects of GSP tariffs on the welfare of recipient countries. According to the model, the extent of this internalization can be observed through a positive relationship between the GSP tariff rate and the share of GSP-eligible goods in the market of the donors countries. Intuitively, this prediction suggests that if donor governments genuinely prioritize welfare of the recipient countries, they are expected to offer deeper GSP

preferences in sectors where firms from the recipient countries have a larger presence.

To empirically test this prediction, we analyze tariff and trade data from all industrialized countries implementing the GSP program between 1988 and 2017. By examining the relationship between GSP tariffs and the market share of GSP-eligible goods, we seek to provide empirical evidence regarding the extent to which GSP donors take into consideration the welfare of recipient countries when determining GSP tariff rates. The main challenge in our empirical strategy is addressing the potential influence of other economic conditions in donor countries on changes in market access and GSP tariff rates. We address this challenge by instrumenting the donor's market share in the recipient country using the recipient country's export share to the rest of the world. The recipient country's export share serves as a strong predictor of its market performance in the donor country, while being exogenous to the donor's GSP policy decision-making process.

The empirical results of our analysis provide support for the causal relationship between the share of GSP-eligible goods in donor market and GSP tariff it choose to apply on these goods. Specifically, our findings reveal that a decrease of one percentage point in the market share of GSP-eligible developing countries is associated with a reduction of 0.08 percentage points in the GSP tariff rate. This result is noteworthy considering the relatively low average GSP preferential tariff rates of 2.37% applied to developing countries and the significant average market share of 25% attained by GSP goods from these countries. Thus, our findings suggest the presence of substantial tariff cooperation that should not be overlooked.

In general, the results indicate that GSP donors do internalize the impact of their GSP tariffs on the profits of firms from developing countries, indicating a certain level of concern for the welfare effects of the GSP program on these countries. However, our empirical analysis does not provide evidence of GSP donors adequately considering the welfare implications of GSP for the least developed countries (LDCs). There are several possible explanations for this lack of evidence in the case of LDCs. These include the initially low tariff rates applied on imports from LDCs (0.72%), the relatively small market share captured by these countries (0.46%), and the challenges highlighted by [Ornelas and](#)

[Ritel \[2020\]](#) regarding infrastructure limitations that hinder LDCs from fully capitalizing on the GSP preferences.

Our findings align with a body of previous studies that have examined the effectiveness of the GSP program. One of the early studies conducted by [Rose \[2004\]](#) found that participation in the GSP program led to a more than 100% increase in trade for an average country. Subsequent studies by [Tomz et al. \[2007\]](#) and [Chang and Lee \[2011\]](#), using a similar empirical strategy based on the gravity model employed by [Rose \[2004\]](#), corroborated the positive aggregate trade impact associated with the GSP status. However, [Subramanian and Wei \[2007\]](#) raised concerns about the identification of GSP status as defined by [Rose \[2004\]](#), highlighting the importance of distinguishing between countries that provide GSP benefits (donors) and countries that receive them (recipients). While Subramanian and Wei also found that GSP donors experience an average increasing imports from recipient countries by over 100%, their results exhibited a negative trade effects in many specific sectors. Moreover, [Herz and Wagner \[2011\]](#), in their dynamic analysis of the GSP program, observed a positive trade effect of GSP in the short term, but a decrease in export values from beneficiaries in the long term. Notably, [Eicher and Henn \[2011\]](#) revealed a completely negative trade effect associated with the GSP program. One possible explanation for this stark difference in findings may be attributed to their broad definition of GSP, which encompasses all other non-reciprocal preferential treatments and the omission of GSPs established in the 1970s.

In addition to examining trade effects, some other studies analyzed GSP efficiency from other perspectives. [Özden and Reinhardt \[2005\]](#) discovered that the GSP status can lead beneficiaries to adopt more protectionist trade policies. For GSP donor countries, [Ketterer et al. \[2015\]](#) emphasized that Japan tend to maintain high preference margin between the Most Favored Nation (MFN) and the GSP rates, which can be viewed as a cost of the GSP program in context of the global trade liberalization. [Hakobyan \[2015\]](#) observed that only 60% of GSP-qualified imports in the US market actually claimed the GSP preference, indicating limited potential of beneficiaries to take advantage of the program. [Manger and Shadlen \[2014\]](#) discovered that a higher export share from GSP recipient countries is associated with an increased likelihood of forming more stable Free Trade Agreements

between GSP donors and recipients.

Overall, the existing literature lacks a definitive consensus regarding the efficiency of the GSP program. The contradictory findings from these previous studies introduce complexity to our understanding of the overall effectiveness of the program. Building upon the insights and complexities revealed by previous studies, our research aims to study the motives of the donor countries behind their design the GSP and improve our understanding of how the program operates.

While previous studies have primarily focused on the external impacts of the GSP status across various dimensions, our study adopts a different approach. We investigate the motive of GSP donors in implementing the GSP program. Specifically, we investigate whether GSP donor countries genuinely consider the economic development of the GSP recipient countries when designing their GSP tariff schemes. To the best of my knowledge, we are the first and only study looking into the motives behind the GSP scheme design by donors. By exploring this aspect of the program, we are able to provide insights into the extent to which the GSP program is effectively aligns with its intended goals.

Our empirical framework relies on a comprehensive identification strategy. As we discussed above, [Subramanian and Wei \[2007\]](#) highlighted the importance of distinguishing between GSP granting countries and GSP-eligible recipients. However, we further differentiate GSP-eligible developing countries from LDCs. This distinction allows us to uncover potentially divergent results for these two groups. Moreover, we address the issue of redundant identification of GSP status caused by the presence of multiple preferential trade regimes available for GSP beneficiaries. We consider a GSP tariff as applicable only if it represents the lowest tariff among all other available preferential tariffs. This allows for a more comprehensive analysis of the effectiveness of the GSP program, accounting for the intricate dynamics of preferential trade agreements. Through our rigorous analysis, we hope to shed light on the underlying puzzle and contribute to the ongoing dialogue surrounding the efficiency of the GSP program in promoting the economic interests of developing countries.

The remaining sections of the paper are structured as follows. Section [2.2](#) provides

an introduction to the theoretical framework that determines the factors influencing the GSP tariff rate. Section 2.3 presents the empirical implementation, including our empirical specification (section 2.3.1), a discussion on limitation (2.3.2) and endogeneity concerns (section 2.3.3) and details about the the data sources used(section 2.3.4). The baseline results are demonstrated in section 2.4, while section 2.5 presents the robustness tests. Finally, section 2.6 provides a concise summary of our findings.

2.2 Theoretical Framework

2.2.1 Theoretical Setup

To analyze whether the concern for economic development in the countries receiving Generalized System of Preferences (GSP) benefits is a driving force behind the GSP policy, it is necessary to understand the factors influencing the choice of GSP preferential tariffs by the donor countries. In this study, we consider an economic model that categorizes all countries into three groups: the GSP donor countries (indexed by $d \in \{1, \dots, D\}$), the GSP recipient countries (indexed by $r \in \{1, \dots, R\}$), and the foreign countries in the rest of the world (ROW) (indexed by $f \in \{1, \dots, F\}$).

Within each country c , including the donor, recipient, and foreign countries, there are $I + 1$ industries (indexed by $i \in \{1, \dots, I\}$). The first industry ($i = 0$) produces the numeraire good, which is assumed to be produced solely using labor under constant returns to scale, setting the wage rate and normalizing the income to one in a competitive equilibrium. In the remaining industries, there are n_{ic} symmetric firms producing and selling good i in the internationally competitive market characterized by monopolistic competition.

The objective of this project is to examine whether GSP donor countries take into account the welfare implications of GSP rates on the recipient countries. From the perspective of consumers in each GSP donor country, goods produced within their own country are considered domestic, while goods produced elsewhere are considered imports. For the sake of tractability, each GSP donor country imports from its GSP-eligible countries and treats

other donor countries the same as foreign countries. As a result, individuals in each GSP donor country are faced with a maximum of $n_{iD} + \sum_{r=1}^R n_{ir} + \sum_{f=1}^F n_{if}$ varieties of good i .

Each individual in this economy maximizes the Cobb-Douglas utility function:

$$U = X_0 + \sum_i^I a_i \ln X_i, \quad \sum_i^I a_i = 1, \quad (2.1)$$

where a_i is the proportion of the total expenditure that each individual spends on goods i . X_0 is the consumer's consumption of the numeraire good and X_i is the aggregate consumption of non-numeraire goods taking the constant elasticity of substitution (CES) form given by

$$X_i = \left(n_{iD} d_{iD}^{\frac{1}{\sigma_i}} x_{iD}^{\frac{\sigma_i-1}{\sigma_i}} + \sum_{r=1}^R n_{ir} d_{ir}^{\frac{1}{\sigma_i}} x_{ir}^{\frac{\sigma_i-1}{\sigma_i}} + \sum_{f=1}^F n_{if} d_{if}^{\frac{1}{\sigma_i}} x_{if}^{\frac{\sigma_i-1}{\sigma_i}} \right)^{\frac{\sigma_i}{\sigma_i-1}}, \quad (2.2)$$

where $\sigma_i > 1$ is the elasticity of substitution between varieties of good i . By allowing consumers have different preferences in the monopolistic competition market, d_{ic} stands for the preference parameter of product i from country c . x_{ic} represents the consumption for product i from country c .

By maximizing utility function (2.1), subject to the standard budget constraint, the demand function can be easily derived as:

$$x_{ic} = a_i d_{ic} p_{ic}^{-\sigma_i} P_i^{\sigma_i-1}, \quad (2.3)$$

where p_{ic} is the price paid by consumers for good i from country c . P_i is the world price also takes the CES form as:

$$P_i = \left(n_{iD} d_{iD} p_{iD}^{1-\sigma_i} + \sum_{r=1}^R n_{ir} d_{ir} p_{ir}^{1-\sigma_i} + \sum_{f=1}^F n_{if} d_{if} p_{if}^{1-\sigma_i} \right)^{\frac{1}{1-\sigma_i}}. \quad (2.4)$$

The government of donor country levies ad-valorem tariff (τ_{ic}) on imported good i originating from country c . Hence, the profits earned by firms in the market of donor countries can be written as:

$$\pi_{ic} = \left(\frac{p_{ic}}{1 + \tau_{ic}} - c_{ic} \right) q_{ic}, \quad (2.5)$$

where c_{ic} represents the marginal cost of firms in country c producing good i , while q_{ic} denotes the quantity of good i supplied by each firm originating from country c in the donor country's market. Considering the assumption that non-numeraire goods are traded in a monopolistic competition market, we fix the world price and derive the optimal price that maximizes the profit function as follows:

$$p_{ic} = \left(\frac{\sigma_i}{\sigma_i - 1} \right) c_{ic} (1 + \tau_{ic}). \quad (2.6)$$

Therefore, after substituting the optimal price (2.6) into the profit function (2.5), we obtain the equilibrium profits:

$$\pi_{ic} = - \frac{p_{ic} q_{ic}}{\sigma_i (1 + \tau_{ic})}. \quad (2.7)$$

In general, donor countries establish two types of tariffs on imported goods. Specifically, the GSP preferential tariff (τ_{iR}) is applied to goods imported from GSP-eligible countries, while the Most Favored Nation (MFN) tariff (τ_{iF}) is applied to goods from other foreign countries in the rest of the world (ROW). It is important to note that τ_{iR} is less than or equal to τ_{iF} , and the tariff rate for imports within the donor country itself, τ_{iD} , is zero.

2.2.2 Trade policy without Tariff Cooperation

If the governments of the donor countries do not prioritize the welfare of recipient countries, they will choose import tariffs that maximize their own national welfare. Therefore, the governments' objective function consists of three components: consumer surplus (CS), producer surplus (PS), and tariff revenue (TR). This objective function can be represented as follows:

$$G(\tau) = CS(\tau) + PS(\tau) + TR(\tau) \quad (2.8)$$

where

$$CS(\tau) = U(X_0, X_i, \tau) - \sum_{i=1}^I \sum_{c=1}^{D+R+F} n_{ic} p_{ic} x_{ic}, \quad (2.9)$$

$$PS(\tau) = \sum_{i=1}^I n_{iD} \pi_{iD} \quad (2.10)$$

and

$$TR(\tau) = \sum_{i=1}^I \left(\frac{\tau_{iR}}{1 + \tau_{iR}} \sum_{r=1}^R n_{ir} p_{ir} x_{ir} + \frac{\tau_{iF}}{1 + \tau_{iF}} \sum_{f=1}^F n_{if} p_{if} x_{if} \right). \quad (2.11)$$

Therefore, the maximization of the government's objective function (2.8) leads to empirically testable predictions for GSP tariffs. The equilibrium GSP preferential tariff (τ_{iR}) can be expressed as follows:

$$\varepsilon_{iR} \frac{\tau_{iR}}{1 + \tau_{iR}} = \frac{\sigma_i - 1}{\sigma_i} s_{iD} + (\sigma_i - 1) \frac{\tau_{iF}}{1 + \tau_{iF}} \sum_{f=1}^F s_{if}, \quad (2.12)$$

where ε_{iR} is the import demand elasticity for goods i exported from recipient countries. The term $s_{ic} = \frac{n_{ic} p_{ic} x_{ic}}{a_i}$ denotes the market share of good i captured by firms in country c within the donor country's market. According to this model, it is clear that the GSP tariff (τ_{iR}) is positively correlated with $\frac{\sigma_i - 1}{\sigma_i} s_{iD}$, which is the first term on the right-hand side of equation (2.15), given the elasticity of substitution (σ_i) is greater than one,. The donor country's market share of domestic products (s_{iD}) is economically linked to the consumer surplus (2.9) in the government's objective function (2.8). A higher value of s_{iD} indicates a greater market share of domestic products, which in turn gives donor governments a stronger incentive to protect their domestic industries through higher GSP tariffs. This reflects the influence of protectionist tendencies in donor countries.

The second term in equation (2.12), $(\sigma_i - 1) \frac{\tau_{iF}}{1 + \tau_{iF}} \sum_{f=1}^F s_{if}$, is derived from the tariff revenue (2.11) component of the government's objective function (2.8). When the donor government lowers the MFN tariff (τ_{iF}), there is a decrease in tariff revenue due to market deviation. To compensate for this loss in tariff revenue, the government needs to reduce the GSP tariff. This relationship between the two types of tariffs is known as

the complementarity effect in the literature. The strength of the complementarity effect depends on the market share of imported goods under the MFN tariff ($\sum_{f=1}^F s_{if}$). If the market share of imported goods under the MFN tariff is larger, the tariff revenue loss will be greater, resulting in a stronger complementarity effect. Therefore, the interaction between the MFN tariff (τ_{iF}) and the market share influenced by the MFN tariff ($\sum_{f=1}^F s_{if}$) is positively correlated with the GSP tariff.

2.2.3 Trade policy with Tariff Cooperation

Nevertheless, if GSP donor countries do exhibit concern for the export-led economic development in their GSP-eligible countries, it is necessary to modify the governments' objective function (2.8) by incorporating the cooperation (CO) term. Consequently, the updated objective function can be expressed as:

$$G(\tau) = CS(\tau) + PS(\tau) + TR(\tau) + \phi CO(\tau) \quad (2.13)$$

where CO, in this project, is the profit of exporting firms from the recipient countries and has the following form:

$$CO(\tau) = \sum_{r=1}^R \sum_{i=1}^I n_{ir} \pi_{ir}. \quad (2.14)$$

The parameter ϕ measures the degree of cooperation between GSP granting countries and their recipient countries. A positive value of ϕ implies that GSP granting countries are actively interested in improving the welfare of their recipient countries. On the other hand, a negative value of ϕ suggests that donor countries prioritize maximizing their own national welfare over the profits of firms from their recipient countries. When there is no additional emphasis placed on tariff cooperation ($\phi = 0$), the final term on the right-hand side of equation (2.13) would be zero. In this case, the updated government objective function (2.13) would be exactly the same as the original objective function (2.8).

Therefore, the maximization of the government's objective function (2.13), taking into account the presence of tariff cooperation, provides empirically testable predictions for GSP

tariffs. The equilibrium GSP preferential tariff (τ_{iR}) can be expressed as follows:

$$\varepsilon_{iR} \frac{\tau_{iR}}{1 + \tau_{iR}} = \frac{\sigma_i - 1}{\sigma_i} s_{iD} + (\sigma_i - 1) \frac{\tau_{iF}}{1 + \tau_{iF}} \sum_{f=1}^F s_{if} - \phi \frac{1}{1 + \tau_{iR}} \left(1 - \frac{\sigma_i - 1}{\sigma_i} \sum_{r=1}^R s_{ir} \right), \quad (2.15)$$

The only difference between equilibrium GSP tariff (2.15) and (2.12) is the third term on the right-hand side of equation (2.15), $-\phi \frac{1}{1 + \tau_{iR}} \left(1 - \frac{\sigma_i - 1}{\sigma_i} \sum_{r=1}^R s_{ir} \right)$, represents the impact of tariff cooperation (2.14) in the government's objective function (2.13). As the elasticity of substitution (σ_i) is greater than 1, $\left(1 - \frac{\sigma_i - 1}{\sigma_i} \sum_{r=1}^R s_{ir} \right)$ is greater than 0. It is important to note that the parameter ϕ measures the extent to which donor countries take into account the economic interests of their GSP-eligible countries when determining the structure of GSP tariff rates. Therefore, $-\phi$ implies that donor governments are more concerned about the economic development of their GSP-eligible countries by reducing the GSP tariff on goods from these countries. Furthermore, the negative correlation between the GSP tariff and the level of cooperation parameter ϕ can be reflected by the positive correlation between the GSP tariff and the donor's market share supplied by GSP-eligible products ($\sum_{r=1}^R s_{ir}$). Intuitively, when recipient countries capture a smaller market share in the donor country, the donor government lowers the GSP preferential tariff to enhance the market competitiveness of goods imported from the recipient countries.

2.3 Empirical Implementation

2.3.1 Empirical Specification

This paper aims to examine whether GSP donor governments internalize the impact of their GSP tariff on the welfare of their GSP-eligible countries. As discussed in the previous section, the theoretical model (2.15) suggests an inversely proportional relationship between the level of internalization (ϕ) and the GSP tariff. However, measuring the parameter ϕ directly is challenging in the existing literature. Therefore, in this study, we empirically estimate the causal relationship between the GSP tariff and the market share captured by

goods from GSP receiving countries in the donor market.

To empirically investigate this relationship, we introduce an additive error term and formulate the empirical specification based on the theoretical model (2.15). The empirical specification can be expressed as follows:

$$\tau_{cit}^{gsp} = \beta_0 + \beta_1 s_{cit}^{mfn} \tau_{cit}^{mfn} + \beta_2 s_{cit}^{gsp} + \alpha_{ci} + \alpha_{it} + \alpha_{ct} + \varepsilon_{cit}, \quad (2.16)$$

where the subscript c is the index for GSP donor countries since GSP preferences are provided only by donor countries to receiving countries. The subscript i indicates the type of product, and the subscript t represents the year. The dependent variable, τ_{cit}^{gsp} , represents the GSP tariff imposed by donor country c on good i imported from its GSP receiving countries in year t . It should be noted that GSP donor countries may apply different GSP tariff rates for eligible developing countries and eligible least developed countries (LDCs). Therefore, we measure τ_{cit}^{gsp} separately for developing countries and LDCs. τ_{cit}^{mfn} represents the MFN tariff imposed by donor country c on imported good i from MFN-eligible countries. s_{cit}^{mfn} denotes the market share captured by imported good i under donor country c 's MFN tariff. s_{cit}^{gsp} represents the market share of good i imported from donor country c 's GSP receiving countries. To account for the distinction between GSP-eligible developing countries and LDCs, we measure s_{cit}^{gsp} separately for these categories. To control for potential unobserved heterogeneity, we include fixed effects. α_{ci} captures country-industry fixed effects, α_{it} captures industry-year fixed effects, and α_{ct} captures country-year fixed effects. Finally, ε_{cit} represents the error term accounting for unobserved factors and random disturbances. The empirical model (2.16) is built upon the theoretical model (2.15). However, due to limitations in the available data, certain adjustments have been made. As a result, it is important to carefully examine the expected signs and significance of the coefficients and their implications, once the discrepancies have been explained in the following section.

2.3.2 Limitation and Implication of Empirical Model

Indeed, the empirical specification (2.16) does not include the elasticity terms present in the theoretical model (2.15). Estimating reliable elasticity values is a challenging task acknowledged in the literature. However, the main prediction of the theoretical model (2.15) is that the GSP tariff is positively associated with the donor's market share of goods from GSP recipient countries. If this causal relationship can be confirmed empirically, it would not be surprising to find that the relationship is stronger for products with larger elasticities. Intuitively, when the elasticity of a particular product is higher, changes in the GSP tariff may have a greater impact on the market share of that product. Therefore, investigating the relationship between the GSP tariff and the donor's market share of goods from GSP recipient countries provides valuable insights, even without explicitly estimating the elasticity terms.

The empirical specification (2.16) also excludes the market share of the donor's domestic products. Although the trade and tariff data used in this study are classified at a highly detailed 6-digit HS product level, obtaining industry output data necessary for calculating the market share of domestic goods is limited to a few developed countries at the 4-digit ISIC level. Consequently, aggregating the product classification would lead to a loss of over 90 percent of the observations. In the robustness section of the paper, we addressed the limitation regarding the market share of domestic products by aggregating the data at the ISIC 4-digit product level. Through this analysis, we found that the market share of the donor's domestic products does not have a significant impact on the GSP tariff towards both GSP-eligible developing countries and LDCs. Thus, it can be concluded that the market share of domestic products does not play a significant role in influencing preferential trade policy decisions in this context.

To overcome the data limitation mentioned earlier, we employ the donor's import value from GSP-eligible countries and other foreign countries as proxies for the market share of GSP-eligible countries (τ_{cit}^{gsp}) and other foreign countries (s_{cit}^{mfn}). Specifically, we calculate the ratio of the import value from GSP-eligible countries and other foreign countries to the total import value. A higher import share in this context indicates a larger market share of

a particular product. By using these proxies, we can approximate and analyze the market shares of GSP-eligible countries and other foreign countries, respectively.

After clarifying all the variables in the empirical model (2.16), we can discuss the expected signs and significance of the coefficients. First, based on the well-established tariff complementarity effect in the literature, we anticipate that the coefficient β_1 should be positive and statistically significant. This effect suggests that as the MFN tariff decreases, the GSP tariff decreases, reflecting the trade-off between tariff revenue and market deviation. Second, the coefficient of interest in this project is β_2 , which captures the relationship between the GSP tariff and the market share of goods from GSP recipient countries. It is expected that β_2 will be positive and statistically significant, supporting the presence of tariff cooperation between GSP donors and their target countries. This implies that GSP donor countries show concern for the welfare of their recipient countries by reducing the GSP tariff to stimulate the market performance of imported goods from GSP recipient countries. By examining the signs and significance of these coefficients, we can gain insights into the factors influencing GSP tariff decisions and the extent to which GSP donors internalize the impact of their tariff policies on the welfare of their recipient countries.

2.3.3 Endogeneity Issues

This section aims to introduce and address the endogeneity problem inherent in the empirical model (2.16). We will first provide an overview of each potential endogeneity issue arising from the donor's import shares from other countries. Subsequently, we will discuss the steps taken to mitigate these concerns and ensure the validity of our empirical analysis.

The empirical model (2.16) is built upon the theoretical model (2.15). However, the presence of the donor's import shares from other countries introduces potential endogeneity issues. This is because preference-granting countries periodically revise their GSP programs, which include determining GSP tariff rates, preference-eligible countries, and preference-eligible industries. The public dissemination of this information can potentially result in reverse causality, where trade partners' export decisions or export volumes to the donor market are influenced.

For instance, if the US Congress decides to reduce the GSP tariff on certain goods from target countries, the anticipation of this decision may prompt those target countries to increase their export volumes to the US market. Similarly, other trade partners not included in the GSP program may recognize that they may lose price competitiveness in the US market and subsequently shift their export focus to other markets.

These endogeneity concerns highlight the need to carefully address the potential bidirectional relationship between GSP tariff rates and market shares. Failure to account for endogeneity may lead to biased and inconsistent estimation results. Therefore, appropriate econometric techniques and strategies must be employed to mitigate endogeneity issues in the analysis.

To address the endogeneity concerns in the empirical model, we employ the two-stage least squares (2SLS) estimator. This approach allows us to obtain consistent estimates by using instrumental variables. In our case, the instrumental variable used to instrument a GSP donor's import share from its recipient countries is the export share of these recipient countries to the rest of the world. This instrument is exogenous because the export share of GSP receiving countries to the rest of the world is unrelated to the donor's GSP policy. However, a higher export share of these GSP receiving countries indicates their comparative advantages in a certain product to some extent. It is highly plausible that these recipient countries would also increase their export values to the donor country that provides trade preferences to them.

Similarly, a donor's import share from other foreign countries not involved in the GSP program is instrumented by the export share of these other foreign countries to the rest of the world. This instrument helps address the endogeneity issue by providing an exogenous variable that is unrelated to the GSP policy.

By using these instrumental variables, we can mitigate the endogeneity problems arising from the potential reverse causality between GSP tariff rates and market shares. The 2SLS estimator enables us to obtain consistent estimates and establish more reliable causal relationships in our analysis.

2.3.4 Data

To conduct a comprehensive analysis of the GSP program, this paper focuses on all industrialized countries that provide GSP preferences to both developing countries and least developed countries (LDCs). The information regarding the GSP-granting countries is obtained from the World Trade Organization (WTO) website. The list of eleven GSP-granting countries included in our sample is presented in table 2.1.

It should be noted that industrialized countries and emerging economies such as Chile, China, and South Korea are not included in our sample, as they only provide GSP preferences to LDCs. The analysis covers the period from 1988 to 2017. Consequently, countries like Armenia and Montenegro are excluded from the list since they implemented their GSP programs after 2016.

Table 2.1: GSP Donor Country List

GSP Provider	Entry Year
Australia	1974
Canada	1974
EU	1971
Iceland	2002
Japan	1971
New Zealand	1972
Norway	1971
Russia	2010
Switzerland	1972
Turkey	2002
USA	1976

The GSP preferential tariff, MFN tariff, and international trade data utilized in this study are collected from the World Integrated Trade Solution (WITS) platform. These data are available at the 6-digit Harmonized System (HS) product classification level. Additionally, the WITS database provides information on GSP-eligible developing countries and least developed countries (LDCs) for each GSP donor country.

To ensure the identification of causality between the GSP tariff and the export performance of the target countries, we specifically isolate the GSP effect from other non-reciprocal

trade preferences. This is accomplished by considering a GSP recipient country as eligible only if its exports to the GSP-granting country are actually subject to the GSP tariff. In cases where a GSP recipient country has also formed a free trade agreement with its GSP donor, the country is removed from the donor’s list of GSP-eligible countries. The preferential trade agreement data used in this process is sourced from the World Trade Organization (WTO).

For industry-level production data pertaining to each GSP donor country, we obtain the information from the United Nations Industrial Development Organization (UNIDO). However, it is important to note that the production data is available only at the 4-digit International Standard Industrial Classification (ISIC) product classification level. This data is utilized to construct the GSP donor country’s market shares of domestic products and imported products from GSP-eligible countries and other foreign countries in the rest of the world. These market shares are used in the robustness testing section.

To ensure consistency and accuracy, we rely on the concordance of country names, country numeric codes, and industry codes obtained from the WITS database. This helps maintain coherence in our data analysis and ensures proper alignment of relevant variables.

The sample used to study the GSP tariff towards developing countries consists of an unbalanced panel with 440,189 observations. The mean values of the corresponding variables in the empirical model (2.16) are presented in the upper part of table 2.2. The average GSP tariff towards eligible developing countries (DC) is 2.37%, while the average MFN tariff is 5.56%. The import share (DC) is calculated as the average of the GSP donor’s total import value from their GSP-eligible developing countries divided by the donor’s total import value. The mean value of import share (DC) is 25%. Similarly, the import share (MFN) is constructed in a similar manner, and the average import share under MFN for a GSP donor country is 40%. Export share (DC) represents the ratio of the total export value from a GSP donor’s GSP-eligible developing countries to the global trade value of a particular product. The mean value of export share (DC) is 29%, which is close to the mean value of import share (DC). Likewise, the average export share (MFN), calculated using a similar method, is 38%, comparable to the mean value of import share (MFN).

The bottom part of Table 2 displays the mean values of variables used to analyze the GSP tariff towards least developed countries (LDCs), with a sample size of 305,380 observations. The average GSP tariff towards eligible LDCs (LDC) is 0.72%, considerably lower than the GSP tariff towards developing countries. The average MFN tariff is 6.65%, slightly different from the MFN tariff presented in the upper part of Table 2 due to the difference in sample composition. The average import share (LDC) is nearly 0.5%, while the average import share (MFN) is 41%, consistent with the import share (MFN) from the upper part of table 2.2. The average export share (LDC) is 1%, and the average export share (MFN) is 40%. These descriptive statistics provide an overview of the average GSP and MFN tariff rates, as well as the import and export shares for both developing countries and LDCs within the sample.

Table 2.2: Summary Statistics

	Variable	Mean	St. Dev.	N
DC	GSP Tariff (DC)	2.37	5.73	440,189
	MFN Tariff	5.56	7.14	440,189
	Import Share (DC)	25.38	30.48	440,189
	Import Share (MFN)	40.18	35.52	440,189
	Export Share (DC)	29.21	26.51	440,189
	Export Share (MFN)	37.88	25.93	440,189
LDC	GSP Tariff (LDC)	0.72	5.43	305,380
	MFN Tariff	6.65	10.24	305,380
	Import Share (LDC)	0.46	3.31	305,380
	Import Share (MFN)	41.13	35.93	305,380
	Export Share (LDC)	0.55	2.58	305,380
	Export Share (MFN)	39.86	28.47	305,380

2.4 Estimation Results

In this section, we present the estimated results for the empirical model (2.16) and discuss their implications for both GSP-eligible developing countries and LDCs.

Table 2.3 presents the estimation results for the empirical specification (2.16) regarding different groups of target countries. In column (1), which focuses on developing countries,

the OLS estimate for β_1 is 0.003, indicating a positive and significant relationship between the GSP tariff and the market share captured by developing countries. The OLS estimate for β_2 is 0.018, also positive and significant as expected, suggesting that GSP donors consider the effect of the GSP tariff on the welfare of GSP-eligible developing countries. We will not focus on the magnitude and implication for OLS results from column (1) since we cannot reject the possibility of endogeneity problems discussed in section (2.3.3).

In column (2), the market shares from MFN-affected countries ($s_{cit}^{mf^n}$) and GSP-eligible developing countries (s_{cit}^{gsp}) are instrumented using each group of countries' export shares to the rest of the world. The instrumental variable (IV) estimate for β_1 is 0.009, providing further support for the tariff complementarity effect between GSP and MFN tariffs. Specifically, given the mean value of import share from MFN impacted countries ($s_{cit}^{mf^n}$) is 40.18%, 0.009 suggests that a one percentage point reduction in the MFN tariff leads to a 0.36 percentage point reduction in the GSP tariff towards developing countries. The significant and positive estimate of 0.079 for β_2 in this column confirms the main focus of this paper. It provides evidence that GSP donors consider, at least partially, the effect of GSP tariffs on the welfare of GSP-eligible developing countries. Specifically, a one percentage point decrease in the market share of GSP-eligible developing countries leads to a 0.079 percentage point reduction in the GSP tariff rate on average. This suggests that GSP donors lower the GSP tariffs in order to improve the market performance of targeted developing countries in their markets, indicating a concern for the welfare of these countries. These results support the predictions from the theoretical model (2.15) and provide empirical evidence of the relationship between GSP tariffs and the economic development of GSP-eligible developing countries.

Moving to columns (3) and (4) which analyze GSP tariffs towards LDCs, the OLS estimate for β_1 in column (3) is 0.003, indicating the presence of the tariff complementarity effect. However, the estimate for β_2 is 0.014, which is not statistically significant. Therefore, we cannot establish a causal relationship between the GSP tariff and the welfare of GSP-eligible LDCs based on this analysis. To further investigate this relationship, we focus on the IV results from column (4), which provide more robust estimates for GSP tariffs

towards LDCs.

Column (4) presents the IV results for equation (2.16) studying GSP tariff towards LDCs. The construction is the same as column (2), with the market share from GSP-eligible LDCs instrumented by the LDCs' export share to the rest of the world. The significant estimate of 0.006 from the first row in column (4) once again confirms the presence of a complementarity effect between GSP tariffs towards LDCs and MFN tariffs. Specially, on average, one percentage point decrease in MFN tariff leads to a 0.25 percentage point decline in GSP tariff towards LDCs. The value of -0.651 from the third row in column (4) represents the IV estimation for β_2 , which is the main focus of this paper. However, the poor significance level of β_2 suggests that we cannot reject the null hypothesis that the value of -0.651 is different from zero. Therefore, we cannot conclude that GSP donors consider the effect of GSP tariffs on the welfare of GSP-eligible LDCs based on this analysis. The absence of significant results for β_2 in the analysis of GSP tariffs towards LDCs (0.72%) could be attributed to several factors. These include the comparatively lower tariff levels applied to LDCs, the relatively smaller market share captured by LDCs (0.46%), and potential infrastructure limitations that may impede LDCs from fully leveraging the advantages provided by GSP preferences.

Overall, these findings provide empirical evidence supporting the positive relationship between GSP tariffs and the market share of GSP-eligible developing countries. However, the significance of the effect on GSP-eligible LDCs requires further examination due to the lack of statistical significance in the estimated coefficients.

2.5 Robustness

In this section, we present the results of a robustness test using an alternative specification to examine the potential presence of significant missing variable bias in our estimation.

As discussed in section (2.3.1), the original empirical model (2.16) was constructed based on the theoretical model (2.15). However, due to data limitations, we were unable to

Table 2.3: Estimation Results

	(1)	(2)	(3)	(4)
	OLS (DC)	IV (DC)	OLS (LDC)	IV (LDC)
$s_{cit}^{mfn} \tau_{cit}^{mfn}(\beta_1)$	0.003*** (3.98)	0.009*** (7.20)	0.003*** (3.84)	0.006*** (4.84)
$s_{cit}^{gsp}(\text{DC})(\beta_2)$	0.018** (2.75)	0.079*** (3.94)		
$s_{cit}^{gsp}(\text{LDC})(\beta_2)$			0.014 (1.15)	-0.651 (-0.41)
$Adj.R^2$	0.8781		0.6840	
N	440,189	440,189	305,380	305,380

Notes: The dependent variable for columns (1) and (2) is τ_{dc}^{gsp} . The dependent variable for columns (3) and (4) is τ_{ldc}^{gsp} . * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country-industry, country-year and industry-year fixed effects and are double clustered at country and industry levels.

include the GSP donor's market share of domestic goods in the empirical specification. It is possible that changes in the GSP tariff may be driven by variations in the market share of domestic goods or that there may be other factors at play that weaken the causality between the GSP tariff and the market share of GSP recipient countries.

To address this concern, we collected industry-level output data from the UNIDO database. Since the available output data is recorded at a 4-digit ISIC product classification, we aggregated the 6-digit HS trade and tariff data to the 4-digit ISIC level for our analysis. We then calculated the market share of domestic goods for the donor countries by taking the ratio of industry shipments (net of exports) to total consumption (total shipments minus net exports). Additionally, we reconstructed the market shares of GSP and MFN goods in the 4-digit ISIC classification using the same methodology. Consequently, we updated the empirical specification (2.16) as follows:

$$t_{cit}^{gsp} = \beta_0 + \beta_1 s_{cit}^{mfn} t_{cit}^{mfn} + \beta_2 s_{cit}^{gsp} + \beta_3 s_{cit}^{donor} + \alpha_{ci} + \alpha_{it} + \alpha_{ct} + \varepsilon_{cit}, \quad (2.17)$$

Where t_{cit} is the ratio of τ_{cit} to $(1 + \tau_{cit})$ derived from equation (2.15), and s_{cit}^{donor} represents the market share of domestic goods. By introducing the new variable s_{cit}^{donor} , we aim to assess whether the estimates for β_1 and β_2 remain comparable to the corresponding

Table 2.4: Estimation Results of Theoretical Model

	(1) IV (DC)	(2) IV (LDC)
$s_{cit}^{mfn} \tau_{cit}^{mfn}$	0.008* (2.22)	0.007 (0.23)
$s_{cit}^{gsp}(\text{DC})$	0.010* (2.21)	
$s_{cit}^{gsp}(\text{LDC})$		0.679 (0.24)
s_{cit}^{donor}	0.364 (1.73)	0.298 (0.24)
N	6,472	4,248

Notes: The dependent variable for columns (1) is $\tau_{dc}^{gsp}/(1 + \tau_{dc}^{gsp})$, while for column (2) is $\tau_{ldc}^{gsp}/(1 + \tau_{ldc}^{gsp})$. * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. t-values in parenthesis. All regressions include country-industry, country-year and industry-year fixed effects and are double clustered at country and industry levels.

estimates in Table 2.3. Additionally, we examine whether the inclusion of β_3 significantly eliminates any missing variable bias.

Table 2.4 reports the IV estimation results for specification (2.17). Column (1) focuses on the GSP tariff towards developing countries, while column (2) examines the GSP tariff towards LDCs. The overall findings continue to suggest that GSP donors take into account the impact of GSP tariffs on the economic interests of targeted developing countries. However, we do not find similar evidence for LDCs. Moreover, the inclusion of the market share of domestic goods for the donor countries does not have a significant effect on the GSP tariffs. Therefore, our baseline results are validated, as the inclusion of the market share of domestic goods does not significantly affect the estimated coefficients for the GSP tariffs towards developing countries and LDCs.

2.6 Conclusion

This study contributes to the understanding of the effectiveness of the Generalized System of Preferences (GSP) program by examining the factors that influence the determination of GSP tariff rates and the extent to which GSP donor countries consider the welfare of recipient countries.

Our theoretical model and empirical analysis provide evidence that GSP donor countries demonstrate a certain level of concern for the welfare of developing countries. The reduction in GSP tariff rates in response to a decline in the market share of GSP-eligible products from developing countries suggests that GSP donors actively lower tariffs to enhance the market performance of these products and promote economic development in recipient countries.

However, our findings also indicate that a similar causal relationship is not observed for GSP-eligible least developed countries (LDCs). This suggests that GSP donors may not prioritize the welfare of LDCs to the same extent as developing countries.

Despite the availability of other non-GSP preferential programs for developing countries, the GSP program remains the most extensive and covers a larger number of beneficiaries. Therefore, further attention and exploration of the GSP field in future academic literature are warranted to gain a comprehensive understanding of its effectiveness and implications.

In summary, this study sheds light on the factors influencing GSP tariff rates and provides valuable insights into the motivations and considerations of GSP donor countries in their trade policies. By examining the relationship between GSP tariffs and market shares of GSP-eligible products, we contribute to the ongoing dialogue on the effectiveness of the GSP program and its impact on the economic development of recipient countries.

References

- O. Al-Ubaydli and P. A. McLaughlin. RegData: A numerical database on industry-specific regulations for all United States industries and federal regulations, 1997–2012. *Regulation & Governance*, 11(1):109–123, March 2017. doi: 10.1111/rego.12107. URL <https://ideas.repec.org/a/wly/reggov/v11y2017i1p109-123.html>.
- K. Bagwell and R. W. Staiger. Regionalism and multilateral tariff co-operation. In *International trade policy and the pacific rim*, pages 157–190. Springer, 1999.
- J. B. Bailey and D. W. Thomas. Regulating away competition: the effect of regulation on entrepreneurship and employment. *Journal of Regulatory Economics*, 52(3):237–254, December 2017. doi: 10.1007/s11149-017-9343-9.
- J. B. Bailey, D. W. Thomas, and J. R. Anderson. Regressive effects of regulation on wages. *Public Choice*, 180(1/2):pp. 91–103, 2019. ISSN 00485829, 15737101. URL <https://www.jstor.org/stable/48704155>.
- D. Chambers and C. O’Reilly. Regulation and income inequality in the United States. *European Journal of Political Economy*, 72(C), 2022. doi: 10.1016/j.ejpoleco.2021.1. URL <https://ideas.repec.org/a/eee/poleco/v72y2022ics0176268021000872.html>.
- D. Chambers, C. A. Collins, and A. Krause. How do federal regulations affect consumer prices? an analysis of the regressive effects of regulation. *Public Choice*, 180:57–90, 2019.
- P.-L. Chang and M.-J. Lee. The wto trade effect. *Journal of International Economics*, 85(1):53–71, 2011.

- B. Coffey, P. A. McLaughlin, and P. Peretto. The cumulative cost of regulations. *Review of Economic Dynamics*, 38:1–21, 2020. ISSN 1094-2025. doi: <https://doi.org/10.1016/j.red.2020.03.004>. URL <https://www.sciencedirect.com/science/article/pii/S1094202520300223>.
- T. S. Eicher and C. Henn. In search of wto trade effects: Preferential trade agreements promote trade strongly, but unevenly. *Journal of International Economics*, 83(2):137–153, 2011.
- A. Estevadeordal, C. Freund, and E. Ornelas. Does regionalism affect trade liberalization toward nonmembers? *The Quarterly Journal of Economics*, 123(4):1531–1575, 2008.
- S. Hakobyan. Accounting for underutilization of trade preference programs: The us generalized system of preferences. *Canadian Journal of Economics/Revue canadienne d'économique*, 48(2):408–436, 2015.
- B. Herz and M. Wagner. The dark side of the generalized system of preferences. *Review of International Economics*, 19(4):763–775, 2011.
- B. Karacavaali and N. Limão. The clash of liberalizations: preferential vs. multilateral trade liberalization in the european union. *Journal of international Economics*, 74(2):299–327, 2008.
- T. D. Ketterer, D. Bernhofen, and C. Milner. Preferences, rent destruction and multilateral liberalization: The building block effect of cusfta. *Journal of International Economics*, 92(1):63–77, 2014.
- T. D. Ketterer, D. M. Bernhofen, and C. Milner. The impact of trade preferences on multilateral tariff cuts: Evidence for japan. *Journal of the Japanese and International Economies*, 38:31–51, 2015.
- D. J. Kuenzel, R. R. Sharma, et al. *Preferential Trade Agreements and MFN Tariffs: Global Evidence*. Wesleyan University, Department of Economics, Public Affairs Center, 2019.

- N. Limão. Preferential trade agreements as stumbling blocks for multilateral trade liberalization: Evidence for the united states. *American Economic Review*, 96(3):896–914, 2006.
- D. S. Lucas and C. J. Boudreaux. National regulation, state-level policy, and local job creation in the united states: A multilevel perspective. *Research Policy*, 49(4):103952, 2020. ISSN 0048-7333. doi: <https://doi.org/10.1016/j.respol.2020.103952>. URL <https://www.sciencedirect.com/science/article/pii/S0048733320300329>.
- J. Mai and A. Stoyanov. The effect of the canada-us free trade agreement on canadian multilateral trade liberalization. *Canadian Journal of Economics/Revue canadienne d'économique*, 48(3):1067–1098, 2015.
- M. S. Manger and K. C. Shadlen. Political trade dependence and north–south trade agreements. *International studies quarterly*, 58(1):79–91, 2014.
- E. Ornelas. Rent destruction and the political viability of free trade agreements. *The Quarterly Journal of Economics*, 120(4):1475–1506, 2005.
- E. Ornelas and M. Ritel. The not-so-generalised effects of the generalized system of preferences. *The World Economy*, 43(7):1809–1840, 2020.
- Ç. Özden and E. Reinhardt. The perversity of preferences: Gsp and developing country trade policies, 1976–2000. *Journal of Development Economics*, 78(1):1–21, 2005.
- A. Postigo. Explaining utilization of free trade agreements by sectoral interests and binding of unilateral concessions. *International Development*, 148:1–63, 2013.
- A. K. Rose. Do we really know that the wto increases trade? *American Economic Review*, 94(1):98–114, March 2004. doi: 10.1257/000282804322970724. URL <https://www.aeaweb.org/articles?id=10.1257/000282804322970724>.
- A. Subramanian and S.-J. Wei. The wto promotes trade, strongly but unevenly. *Journal of international Economics*, 72(1):151–175, 2007.

M. Tomz, J. L. Goldstein, and D. Rivers. Do we really know that the wto increases trade?
comment. *American Economic Review*, 97(5):2005–2018, 2007.

Chapter 3

The Effect of Regulation on the Employment Rate in the U.S.

3.1 Introduction

Our project aims to contribute to the existing literature by examining the impact of regulation on the local labor market in the United States. The increasing intensity of national regulations in developed countries, including the United States, Canada, Britain, and the European Union, is a prominent trend that cannot be ignored. This is evident in the significant growth of the Federal Regulation code in the U.S., which has nearly tripled in length over the past five decades. Similar patterns of regulatory expansion can be observed in other countries, as governments extend their reach into emerging areas such as climate and data protection. This expansion has sparked a spirited debate between proponents of the public interest, who argue that regulations address market failures and inefficiencies, and advocates of public choice, who emphasize the intangible compliance costs imposed on businesses by regulations. Unfortunately, these hidden costs often go unnoticed by policymakers who propose new regulations, creating a critical oversight in the existing literature.

In light of these considerations, our project aims to comprehensively investigate the

economic consequences of regulations and shed light on their true impact. By analyzing the local labor market, we can provide insights for the U.S. regulatory decision-making process, particularly in conducting proper cost-benefit analyses. Moreover, it is essential to recognize that policymakers cannot be treated as a homogeneous group, as their ideologies and policy approaches vary under different ruling parties. In addition to examining the average effect of regulation, it is equally crucial to explore the relative differences in the cumulative effects introduced by policymakers with diverse political ideologies. This nuanced approach will enable a comprehensive understanding of the multifaceted relationship between labor market outcomes, regulation, and the influence of political parties in shaping the economic landscape of the United States.

To empirically examine the impact of regulation on the local labor market, we construct a panel sample comprising 741 commuting zones over a five-decade period starting from 1970. To ensure the robustness of our analysis and to mitigate the potential influence of the COVID-19 pandemic, we exclude data from the year 2020 and use 2019 as the final year of analysis. Our outcomes of interest, the unemployment rate and gini coefficient at the commuting zone level, serve as indicators of local market vitality. The two indicators are calculated based on the U.S. census data from the Integrated Public Use Microdata Series database. To quantify the intensity of regulation, we utilize the RegData database ([Al-Ubaydli and McLaughlin \[2017\]](#)), which provides the number of restrictive words in each section of the Code of Federal Regulation (CFR) in the U.S. Additionally, the database employs machine learning techniques to analyze the regulatory text of the CFR and assess the applicability of each section to different industries in the U.S. economy. The product of the count of restrictive words and their corresponding industrial relevance ratio represents the industrial restrictiveness of each section of the CFR. We then aggregate the industrial restrictiveness by weighting it with the corresponding employment share in the 4-digit NAICS in the commuting zone level.

To explore the relative influence of regulation introduced by different political parties in the U.S., we interact the decennial cumulative change of regulation with the corresponding political party that issued the regulation. By defining regulations introduced by the

Democratic or Republican party based on the party controlling the presidency and the majority of the Senate, we can capture the effects of cumulative regulations implemented under specific ruling parties relative to other political situations. To gather information on the political party affiliation of the president and each senator, we utilize the WebVoteView database.

Our findings reveal that regulation exacerbates local unemployment and income inequality. Specifically, a one standard deviation increase in the number of restrictive words in the CFR, corresponds to an approximate 1.1 percentage point increase in the local unemployment rate. This represents approximately one-fifth of the total unemployment rate. Considering local income inequality, a one standard deviation increase in the number of restrictive words leads to a rise of 2.8 percentage points in the local gini coefficient, which is larger than the one standard deviation of the local gini coefficient. Moreover, the effects of cumulative regulations introduced by Democratic and Republican governments differ significantly from other political situations. On average, a one standard deviation increase in the number of regulatory words introduced by a Republican government over the previous decade is associated with a relative decrease of approximately 0.02 percentage points in the local unemployment rate compared to other political situations. In contrast, a similar increase in the number of regulatory words introduced by a Democratic government leads to a relative increase of approximately 0.01 percentage points in the local unemployment rate. These results hold true when considering the local gini coefficient as the outcome of interest. For an average commuting zone, a one-unit increase in regulatory words introduced by either a Republican or Democratic government is associated with a relative decrease of approximately 0.02 percentage points or an increase of approximately 0.01 percentage points, respectively, compared to other forms of government.

Our results are robust to alternative measures of regulation, as we replace the count of restrictive words with the share of restrictive words over the total words in each CFR section. In all of our specifications, we include commuting zone and year fixed effects to account for unobserved heterogeneity. Furthermore, we include lagged values of the corresponding outcome variables to address potential autoregressive effects. When examining the influence

of cumulative regulation introduced by different ruling parties, we control for regulation levels in both the current and previous periods to capture dynamic effects. Overall, our findings suggest that the costs of regulation, which are often overlooked by policymakers, are indeed reflected in the local labor market. Furthermore, the costs of regulation are even more pronounced when they are implemented by a Democrats-dominated government compared to other forms of government. This finding contrasts with the philosophy of the Democratic party, which aims to promote income equity and job opportunity equality. There are two possible explanations for this discrepancy. Firstly, it is often argued that the Democratic party prioritizes income equality through intense government regulation. However, our baseline results confirm that the intensity of regulation is positively correlated with income inequality. This suggests that the regulations implemented by a Democratic government may have unintended consequences or may not effectively achieve their intended goals in reducing income inequality. Secondly, the Republican party is often associated with prioritizing the growth and expansion of businesses through deregulation. This focus on business-friendly environment may lead to increased employment opportunities.

Overall, our research findings provide further evidence for the negative impacts of government regulations on income equality and employment, aligning with the results of previous studies in the field. [Coffey et al. \[2020\]](#) demonstrated that federal regulations had a significant negative effect on GDP growth, causing a substantial reduction in economic output. Additionally, [Chambers et al. \[2019\]](#) highlighted the contribution of regulations to income inequality, as they disproportionately affected low-income groups through increased consumer prices. [Bailey et al. \[2019\]](#) further emphasized the adverse impact of restrictive regulations on wage growth, affecting different income levels and occupations. Furthermore, government regulations hindered job opportunities, as evidenced by [Lucas and Boudreaux \[2020\]](#) and [Bailey and Thomas \[2017\]](#). These studies revealed a negative relationship between national regulations and net job creation, with regulations also impeding entrepreneurship and the formation of new firms. Our research aligns with these studies, as we find that government regulations have significant implications across various economic activities. Importantly, [Chambers and O'Reilly \[2022\]](#) reinforced the robustness of our findings by

investigating the relationship between federal regulation and income inequality on a state level. Their analysis, although employing a different regulation quantification (FRASE index), showed a similar trend of increasing regulations contributing to rising income inequality. Given these findings, policymakers should carefully consider the potential economic impact of regulations before their implementation. Conducting thorough cost-benefit analyses and considering the potential consequences on income distribution and job opportunities are vital steps in ensuring the effectiveness of regulatory policies. By taking into account the lessons learned from our research and the broader literature, policymakers can make informed decisions that promote economic growth, income equality, and job creation.

This project represents a significant contribution by shifting the analytical focus from traditional country, state, or county-level analysis to a more detailed and contextually relevant level—the commuting zone. This approach is particularly important considering that many policy decisions and interventions are implemented at the regional or metropolitan level. By examining the economic impact of regulation within commuting zones, our research offers insights for policymakers and practitioners engaged in regional development and planning. Moreover, our study introduces a pioneering perspective by disaggregating the effects of regulation based on the political party affiliation of policymakers. While previous studies have explored the economic consequences of regulation across different levels, they often treated policymakers as a homogeneous group, neglecting potential variations in regulatory outcomes associated with political ideologies. In contrast, our research takes a novel approach by explicitly distinguishing the cumulative effects of regulations introduced by Democratic and Republican governments. This unique perspective enables a deeper understanding of how political ideologies shape the economic consequences of regulation.

The remainder of this paper is structured as follows. In section 3.2, we describe our data (section 3.2.1) and empirical models (section 3.2.2 and section 3.2.3). Section 3.3 exhibits the empirical results and their implications. Section 3.4 concludes.

3.2 Empirical Specification

3.2.1 Data

To examine the impact of regulation on economic activities, we construct a panel dataset using commuting zones (CZs) over five decennial census years spanning from 1970 to 2019 in the United States. Our sample consists of all 741 CZs in the U.S. classified in year of 1990.

We measure the main dependent variables, namely the unemployment rate and income inequality (GINI Index), at the CZ level. These variables are calculated using data from the Integrated Public Use Microdata Series (IPUMS-USA), which provides U.S. decennial census data with information on counties and industries in which each resident resides. County and industry information is classified using temporal classification systems specific to the census year. To capture labor market dynamics and regional policies, we convert county-level data into commuting zone level data using the methodology provided by David Autor and David Dorn (2013). Different industrial classifications are reconciled using the 4-digit North American Industry Classification System (NAICS) codes, which are available from the United States Census Bureau.

To assess the extent of regulation, we utilize the RegData database from the Mercatus Center at George Mason University. This database provides information on the total number of words and the number of restrictive words (e.g., 'shall,' 'must,' 'may not,' 'required,' and 'prohibited') in each part of the Code of Federal Regulations (CFR). Additionally, the database offers the applicability of each CFR to specific industries at the 2-, 3-, and 4-digit NAICS levels in the U.S. economy, expressed as a ratio between 0 and 1. We quantify industry-specific federal regulation by summing the number of restrictive words in each CFR multiplied by its corresponding industrial relevance ratio for each year. To capture CZ-level regulatory intensity, we aggregate federal regulation by weighting it with the corresponding employment share of each industry in a CZ for a given year. To facilitate interpretation and normalize the variable range, we divide the weighted number of restrictive words by 10,000. Furthermore, we construct an alternative measure of regulation by replacing the

number of restrictive words with the ratio of restrictive words to total words in each CFR, addressing variations in text length.

The RegData database also provides information on the department or agency responsible for publishing each CFR, allowing us to explore the impact of regulatory costs when the governing party (Democratic or Republican) is different. Since the heads of executive branch departments are typically appointed by the president and confirmed by the Senate, we construct a Democratic dummy variable with a value of one when both the president and the Senate majority are Democrats, and a Republican dummy variable when both are Republicans. We multiply each dummy variable by the annual change in restrictive words and aggregate this product for each decade to analyze the additional effects of regulations issued by each party on economic activities. Information on the political party affiliation of the president and each senator is collected from the WebVoteView database operated by the UCLA Political Science Division.

Table 3.1 provides a summary of the key variables used in the analysis. The average unemployment rate at the commuting zone level is 6.5%, calculated as the ratio of unemployed individuals to the labor force within each CZ, ranging from 0.014 to 0.224. Income inequality, as measured by the Gini coefficient, has an average value of 0.432, ranging from 0.362 to 0.552 across CZs. To capture the extent of regulation, we employ two measures: the industry-weighted count of restrictive words and the industry-weighted share of restrictive words within each commuting zone. We present the one-period lagged value for both measures. The normalized number of restrictive words, on average, is 0.022, with a range of 0.004 to 0.062. Similarly, the average industry-weighted share of restrictive words is 0.017, ranging from 0.004 to 0.044. Considering the political dynamics, the average decennial change in the number of restrictive words issued by the Democratic party is 0.004, while the corresponding figure for the industry-weighted share of restrictive words is 0.001. When interacting these measures with the Republican dummy variable, the mean values are 0.002 and 0.001, respectively.

Table 3.1: Summary Statistics

Variable	Mean	St. Dev.	Min	Max	N
Unemployment Rate	0.065	0.030	0.014	0.224	3,705
Gini Index	0.432	0.024	0.362	0.552	3,705
l.Restriction_word	0.022	0.011	0.004	0.062	3,705
l.Restriction_share	0.017	0.008	0.004	0.044	3,705
d.Restriction_word_D	0.004	0.004	-0.003	0.023	3,705
d.Restriction_share_D	0.001	0.002	-0.004	0.011	3,705
d.Restriction_word_R	0.002	0.002	-0.002	0.008	3,705
d.Restriction_share_R	0.001	0.002	-0.003	0.008	3,705

3.2.2 Empirical Model

To empirically examine our research question, we employ Ordinary Least Squares (OLS) with fixed effects to estimate the effects of regulation on the unemployment rate and Gini coefficient at the commuting zone (CZ) level. The specified model is as follows:

$$Y_{ct} = \beta_0 + \beta_1 l.R_{ct} + \beta_2 l.Y_{ct} + \alpha_c + \alpha_t + \mu_{ct} \quad (3.1)$$

The subscript c represents commuting zones and t is the index for decennial census years, indicating that our dataset is a longitudinal panel with a decade gap. The dependent variable, denoted as Y_{ct} , corresponds to either the unemployment rate or the Gini coefficient in CZ c in year t . These two indicators are commonly used to assess the economic conditions within a region.

The first independent variable, $l.R_{ct}$, on the right-hand side of equation (3.1), represents

the lagged value of the quantified regulation derived from the Code of Federal Regulations (CFR). We consider two measures for this variable. The first measure is the industry-weighted count of restrictive words in CZ c and year t , while the second measure is the industry-weighted share of restrictive words over the total words in CZ c and year t . By using the lagged value of the regulation, we aim to establish the precedence of regulation from the CFR, ensuring that reverse causality is minimized. As a result, we conduct four separate OLS regressions to estimate the effects of the two regulation measures on the two economic indicators. The coefficient β_1 is of particular interest in this study. In general, if the estimate of β_1 is negative and statistically significant, it intuitively suggests that public administration addresses labor market failures, aligning with the public interest view. Conversely, if β_1 is positive and statistically significant, it indicates that public administration imposes compliance costs that manifest in the labor market, aligning with the public choice view.

The inclusion of the second variable, $l.Y_{ct}$, on the right-hand side of equation (3.1), represents the lagged value of the corresponding dependent variable. By incorporating this control variable, we account for various factors that can influence both the unemployment rate and Gini coefficient, such as economic conditions, policy changes, and seasonal fluctuations. Controlling for past unemployment or Gini coefficient allows us to isolate the specific impact of national regulation on these economic indicators, providing a more accurate understanding of their relationship. The parameters α_c and α_t represent CZ fixed effects and year fixed effects, respectively, capturing unobserved heterogeneity specific to each CZ and time period. Finally, the error term μ_{ct} is assumed to be independently and identically distributed, and we follow the common practice of clustering the standard errors at the CZ level to account for potential correlations within CZs.

3.2.3 Model with Disaggregated Regulation

In addition to examining the general local economic costs of regulatory intensity, we aim to explore whether the impact of regulation from the Code of Federal Regulations (CFR), introduced by different ruling parties, differs. To achieve this, we expand the baseline model

(3.1) by introducing two new variables that capture the cumulative regulations attributed to the Democratic and Republican ruling parties, respectively. The updated model (3.2) can be expressed as follows:

$$Y_{ct} = \gamma_0 + \gamma_1 d.R_{ct}R_t + \gamma_2 d.R_{ct}D_t + \gamma_3 R_{ct} + \gamma_4 l.R_{ct} + \gamma_5 l.Y_{ct} + \alpha_c + \alpha_t + \mu_{ct} \quad (3.2)$$

The subscript notation and the measures of the dependent variable in equation (3.2) remain the same as in equation (3.1). The first independent variable, $d.R_{ct}R_t$, on the right-hand side of equation (2), represents the summation of the annual change in the number of restrictive words multiplied by the Republican dummy variable within a decennial census period. This variable captures the restrictions added over the previous decade during years when both the president and the Senate were controlled by the Republican party. Similarly, the second independent variable, $d.R_{ct}D_t$, is constructed by summing the annual change in the number of restrictive words multiplied by the Democratic dummy variable within the decennial census period. This variable represents the restrictions added over the previous decade during years when both the president and the Senate were controlled by the Democratic party. The coefficients γ_1 and γ_2 are of primary interest in model (3.2). Similar to the baseline model (3.1), both of these variables have alternative measures where the count of restrictive words is replaced by the share of restrictive words. A positive and statistically significant estimate of γ_1 (or γ_2) would indicate that when the government (president and majority senators) operates entirely under the Republican (Democratic) party, the cumulative regulation disproportionately harms the local labor market compared to other political situations. Conversely, a negative and statistically significant estimate of γ_1 (or γ_2) would suggest that when the government is fully controlled by the Republican (Democratic) party, the cumulative costs of regulation, reflected in the labor market, are relatively less burdensome compared to other situations.

Furthermore, the model includes three control variables, R_{ct} , $l.R_{ct}$ and $l.Y_{ct}$. The first control variable (R_{ct}) represents the level of total restriction in CZ c at the current period t .

It captures the overall regulatory intensity in the specific commuting zone during the current period. The second control variable ($l.R_{ct}$), represents the level of total restriction in CZ c at the previous period, $t-1$. It captures the regulatory intensity in the specific commuting zone during the preceding period, providing information on the dynamic effect of regulation. The last control variable is the one-period lagged value of the corresponding dependent variable, $l.Y_{ct}$. This control variable captures the influence of the previous period's economic indicator on the current period's indicator, helping control for autoregressive effects. By including these control variables, model (3.2) accounts for the autoregressive nature of the dependent variable and captures the dynamic effects of regulatory intensity on the economic indicators. Therefore, coefficient γ_1 represents the additional effect of cumulative regulation introduced by Republican party, given the level of general regulation in the current and previous periods. Similarly, γ_2 represents the additional effect of cumulative regulation introduced by the Democratic party. The parameters α_c and α_t capture CZ fixed effect and year fixed effect, respectively, accounting for unobserved heterogeneity specific to each CZ and time period. The error term μ_{ct} is assumed to be independently and identically distributed. As with the baseline model, we follow the common practice of clustering the standard errors at the CZ level to address potential correlations within CZs.

3.3 Empirical Results

Table 3.2 reports the estimated result for our baseline model (3.1). The table includes commuting zone and year fixed effects, and the standard errors are clustered at the commuting zone level. The dependent variable is the unemployment rate in commuting zone c and year t in the first four columns, and the gini index in commuting zone c and year t in the last four columns.

In columns (1) and (2), we investigate the relationship between local unemployment and the one-period lagged count of regulatory words, with and without controlling for previous unemployment, respectively. In both columns, we observe a positive relationship between local unemployment and the count of regulatory words. However, the adjusted R-squared

value of the model improves from 0.7181 in column (1) to 0.7208 in column (2) when we include the control variable of previous unemployment. This improvement suggests that controlling for previous unemployment helps to mitigate the influence of unobserved factors on the current unemployment rate. To ensure a more accurate estimation of the relationship between local unemployment and the count of regulatory words, we focus on the magnitude of the estimate of β_1 from column (2).

In column (2), the estimated coefficient (β_1) of 0.993 suggests that, on average, a 10,000 increase in regulatory words is associated with a 0.993-unit increase in the local unemployment rate. Given that the standard deviation of the number of restrictive words is 110 (table 3.1), this corresponds to an approximate increase of 1.1 percentage points in the unemployment rate. Combining the mean value of local unemployment rate is 6.5% (table 3.1), around one fifth of the unemployment should be attributed by regulation.

Similarly, in columns (3) and (4), we examine the relationship between the local unemployment rate and the share of regulatory words, with and without the control variable, respectively. As with the count of regulatory words, both columns (3) and (4) show a positive correlation between the local unemployment rate and the intensity of regulation measured by the share of regulatory words. The focus is on the significant estimate of β_1 , which is 1.543 in column (4). This implies that, on average, a one percentage point increase in the share of regulatory words is associated with a 1.543 percentage point increase in the unemployment rate, holding other variables constant.

In columns (5) to (8), we examine the relationship between the gini index and the local regulation intensity, with and without controlling for the previous-period gini index. Similar to the unemployment rate, the gini index measures income inequality on a scale from 0 to 1, where 0 represents perfect equality and 1 represents perfect inequality. It is expected that the coefficients in columns (5) to (8) will move in the same direction as those in columns (1) to (4) that study the unemployment rate. To account for the potential influence of previous-period income inequality, we include the control variable of previous-period gini index in columns (6) and (8). Comparing the adjusted R-squared values between these columns and their corresponding columns without controls (columns (5) and (7)),

we consistently observe higher adjusted R-squared values when the control variables are included. Therefore, we primarily focus on interpreting the estimated coefficients from column (6) and (8) as they provide a more reliable assessment of the relationship between regulatory words and the gini index.

The coefficient of 2.502 in column (6) indicates that, on average, a rise of 10,000 regulatory words is linked to a 2.502-unit increase in the local gini index. This finding implies that if the number of regulatory words were to increase by its standard deviation value of 110, the local gini index would experience a substantial rise of 2.8 percentage points. Notably, this increase is greater than the standard deviation value of local gini index , 0.023 (table 3.1), highlighting the significance of the observed effect. Consistently, the coefficient of 2.439 in column (8) indicates that, on average, a one percentage point increase in the share of regulatory words corresponds to a 2.439 percentage point increase in the local gini index. Overall, the baseline results from table 3.2 confirm that the increase in regulatory intensity is detrimental to local employment and income equity, as evidenced by the positive coefficients in both the unemployment rate and gini index models.

Table 3.2: Estimated Results for the Effect of Regulation on Local Economic Activities

Dep. Var.:	(1) UR	(2) UR	(3) UR	(4) UR	(5) Gini	(6) Gini	(7) Gini	(8) Gini
$l.R_{ct}(\beta_1)$ (count of words)	0.768*** (0.142)	0.993*** (0.153)			3.014*** (0.182)	2.502*** (0.190)		
$l.R_{ct}(\beta_1)$ (share of words)			1.158*** (0.224)	1.543*** (0.236)			3.126*** (0.222)	2.439*** (0.220)
$l.Y_{ct}$		-0.099*** (0.024)		-0.117*** (0.022)		0.163*** (0.025)		0.206*** (0.024)
Adj. R^2	0.7181	0.7208	0.7197	0.7233	0.6491	0.6576	0.6312	0.6456
N	3,705	3,705	3,705	3,705	3,705	3,705	3,705	3,705

Notes: The dependent variable is the unemployment rate in commuting zone c in year t from column (1) to (4). The dependent variable is the gini index in commuting zone c in year t from column (5) to (8). * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. Standard errors are presented in parenthesis and clustered at commuting zone level.. All specifications include commuting zone and year fixed effects.

Table 3.3 presents the estimated results for specification (3.2), which includes the

cumulative regulations attributed to the Republican and Democratic ruling parties. All columns include a full set of fixed effects, and the standard errors are clustered at the commuting zone level. The control variables of the level of regulation at the current period, previous period, and the one-period lag of the corresponding dependent variable are also included.

In columns (1) and (2), the dependent variable is the unemployment rate in commuting zone c and year t . In column (1), the negative and significant estimate of γ_1 (-0.097) indicates that, on average, an increase of 10,000 regulatory words introduced by a complete Republican government over the previous decade leads to a 0.097-unit decrease in the unemployment rate compared to other forms of government. This suggests that if the number of regulatory words added by a complete Republican government were increased by 20 (equivalent to its one standard deviation), the unemployment rate would decrease by approximately 0.02 percentage points compared to other political situations. Conversely, the positive and significant estimate of γ_2 (0.032) shows that an increase of 10,000 regulatory words introduced by a complete Democratic government over the previous decade results in a 0.032-unit increase in the unemployment rate relative to other political situations. Similarly, if the number of regulatory words added by a complete Democratic government were increased by 40 (its one standard deviation), the unemployment rate would increase by approximately 0.01 percentage points compared to other political situations.

In column (2), the measure of local regulation is changed from the number of restrictive words to the share of restrictive words, and consistent results are obtained compared to column (1). The estimated result of γ_1 (-0.158) indicates that a one percentage point increase in the share of restrictive words introduced by a complete Republican government over the previous decade leads to a 0.158 percentage point decrease in the unemployment rate relative to other forms of government. Similarly, the estimated result of γ_2 (0.028) indicates that a one percentage point increase in the share of restrictive words introduced by a complete Democratic government over the previous decade results in a 0.028 percentage point increase in the unemployment rate relative to other forms of government.

In columns (3) and (4) of table 3.3, the dependent variable is the Gini index in commuting

zone c and year t . The independent variables and controls in column (3) are the same as in column (1). The estimated value of γ_1 (-0.087) from column (3) indicates that an one standard deviation increase of regulatory words introduced by a complete Republican government over the previous decade (20) leads to a 0.02 percentage point decrease in the local Gini index compared to other forms of government. Conversely, the estimated value of γ_2 (0.030) indicates that an one standard deviation increase of regulatory words introduced by a complete Democratic government over the previous decade (40) leads to a 0.01 percentage point increase in the local Gini index compared to other forms of government.

In column (4), where the measure of local regulation is the share of restrictive words, similar results are obtained. The estimated value of γ_1 (-0.088) indicates that a one percentage point increase in the share of restrictive words introduced by a complete Republican government over the previous decade leads to a 0.088 unit decrease in the local Gini index relative to other forms of government. Similarly, the estimated value of γ_2 (0.019) indicates that a one percentage point increase in the share of restrictive words introduced by a complete Democratic government over the previous decade leads to a 0.019 unit increase in the local Gini index relative to other forms of government.

Therefore, the results from table 3.3 consistently indicate that regulations introduced by a Republican government are associated with a relative decrease in local unemployment and income inequality, while regulations introduced by a Democratic government are associated with a relative increase in local unemployment and income inequality. This finding is interesting as it contrasts with the common perception that the Democratic Party is often associated with prioritizing social welfare and addressing income inequality, while the Republican Party is often associated with prioritizing business interests and economic growth. There are two possible explanations for this discrepancy. Firstly, while the Democratic party is often seen as prioritizing income equality through stringent government regulation, our baseline results show a positive correlation between regulation intensity and income inequality. This suggests that the regulations implemented by a Democratic government may have unintended consequences or may not effectively achieve their intended goals

of reducing income inequality. Secondly, the Republican party's emphasis on promoting business growth and expansion may contribute to increased employment opportunities.

Table 3.3: Regulation Added by Different Ruling Parties

Dep. Var.:	(1) UR	(2) UR	(3) Gini	(4) Gini
$d.R_{ct}R_t(\gamma_1)$	-0.097***		-0.087***	
(count of words)	(0.006)		(0.007)	
$d.R_{ct}D_t(\gamma_2)$	0.032***		0.030***	
(count of words)	(0.002)		(0.002)	
R_{ct}	3.208***		2.140***	
(count of words)	(0.245)		(0.213)	
$l.R_{ct}$	-1.148**		1.244***	
(count of words)	(0.225)		(0.243)	
$d.R_{ct}R_t(\gamma_1)$		-0.158***		-0.088***
(share of words)		(0.009)		(0.008)
$d.R_{ct}D_t(\gamma_2)$		0.028***		0.019***
(share of words)		(0.002)		(0.001)
R_{ct}		4.405***		2.636***
(share of words)		(0.289)		(0.246)
$l.R_{ct}$		-0.639**		1.353***
(share of words)		(0.286)		(0.260)
$l.Y_{ct}$	-0.061***	-0.090***	0.127***	0.156***
	(0.025)	(0.023)	(0.026)	(0.024)
Adj. R^2	0.7506	0.7622	0.6788	0.6673
N	3,705	3,705	3,705	3,705

Notes: The dependent variable is the unemployment rate in column (1) and (2) and the gini index in in column (3) to (4). * significant at 10% , ** significant at 5% , *** significant at 1% confidence level. Standard errors are presented in parenthesis and clustered at commuting zone level. All specifications include commuting zone and year fixed effects.

3.4 Conclusion

This project attempts to contribute the existing literature by examining the impact of regulation on the local labor market in the United States using novel regulation intensity and census data. Our findings demonstrate that regulations have a detrimental effect on

local unemployment and income inequality, providing insights into the consequences of regulatory policies. The increase in the number of regulatory words is associated with a significant rise in the local unemployment rate and gini coefficient, highlighting the negative impact on both job opportunities and income distribution.

Furthermore, our research stands out by disaggregating the effects of regulation based on the political party affiliation of policymakers. By distinguishing between Democratic and Republican governments, we uncover significant variations in the cumulative effects of regulations. Surprisingly, the accumulation of regulations under Republican governments tends to attenuate the negative effect on local employment and income equity, while regulations under Democratic governments relatively exacerbate the negative impact on the local labor market. These findings challenge the expectations based on the Democratic party's commitment to promoting income equity and job opportunity equality for the entire society. We offer possible explanations, suggesting that the regulations implemented by a Democratic government may fail to effectively achieve their intended goals of reducing income inequality, while the emphasis on creating a business-friendly environment by the Republican party may contribute to increased employment opportunities.

In conclusion, our project provides insights into the relationship between regulation, political ideologies, and economic outcomes within the context of the local labor market. These findings have important implications for public policy, emphasizing the need for comprehensive cost-benefit analyses to inform regulatory decision-making and ensure the effectiveness and efficiency of policies. Unfortunately, the lack of standardized practices in assessing the cost of regulation after implementation poses a challenge in determining whether the benefits of a policy justify the costs. Therefore, our research underscores the importance of conducting thorough cost-benefit analyses to inform policy discussions and promote sustainable economic development and equitable outcomes at the regional level. The knowledge generated from this study provides a foundation for further exploration and informs policy discussions aimed at fostering inclusive and prosperous societies.

References

- O. Al-Ubaydli and P. A. McLaughlin. RegData: A numerical database on industry-specific regulations for all United States industries and federal regulations, 1997–2012. *Regulation & Governance*, 11(1):109–123, March 2017. doi: 10.1111/rego.12107. URL <https://ideas.repec.org/a/wly/reggov/v11y2017i1p109-123.html>.
- K. Bagwell and R. W. Staiger. Regionalism and multilateral tariff co-operation. In *International trade policy and the pacific rim*, pages 157–190. Springer, 1999.
- J. B. Bailey and D. W. Thomas. Regulating away competition: the effect of regulation on entrepreneurship and employment. *Journal of Regulatory Economics*, 52(3):237–254, December 2017. doi: 10.1007/s11149-017-9343-9.
- J. B. Bailey, D. W. Thomas, and J. R. Anderson. Regressive effects of regulation on wages. *Public Choice*, 180(1/2):pp. 91–103, 2019. ISSN 00485829, 15737101. URL <https://www.jstor.org/stable/48704155>.
- D. Chambers and C. O’Reilly. Regulation and income inequality in the United States. *European Journal of Political Economy*, 72(C), 2022. doi: 10.1016/j.ejpoleco.2021.1. URL <https://ideas.repec.org/a/eee/poleco/v72y2022ics0176268021000872.html>.
- D. Chambers, C. A. Collins, and A. Krause. How do federal regulations affect consumer prices? an analysis of the regressive effects of regulation. *Public Choice*, 180:57–90, 2019.
- P.-L. Chang and M.-J. Lee. The wto trade effect. *Journal of International Economics*, 85(1):53–71, 2011.

- B. Coffey, P. A. McLaughlin, and P. Peretto. The cumulative cost of regulations. *Review of Economic Dynamics*, 38:1–21, 2020. ISSN 1094-2025. doi: <https://doi.org/10.1016/j.red.2020.03.004>. URL <https://www.sciencedirect.com/science/article/pii/S1094202520300223>.
- T. S. Eicher and C. Henn. In search of wto trade effects: Preferential trade agreements promote trade strongly, but unevenly. *Journal of International Economics*, 83(2):137–153, 2011.
- A. Estevadeordal, C. Freund, and E. Ornelas. Does regionalism affect trade liberalization toward nonmembers? *The Quarterly Journal of Economics*, 123(4):1531–1575, 2008.
- S. Hakobyan. Accounting for underutilization of trade preference programs: The us generalized system of preferences. *Canadian Journal of Economics/Revue canadienne d'économique*, 48(2):408–436, 2015.
- B. Herz and M. Wagner. The dark side of the generalized system of preferences. *Review of International Economics*, 19(4):763–775, 2011.
- B. Karacavaali and N. Limão. The clash of liberalizations: preferential vs. multilateral trade liberalization in the european union. *Journal of international Economics*, 74(2):299–327, 2008.
- T. D. Ketterer, D. Bernhofen, and C. Milner. Preferences, rent destruction and multilateral liberalization: The building block effect of cusfta. *Journal of International Economics*, 92(1):63–77, 2014.
- T. D. Ketterer, D. M. Bernhofen, and C. Milner. The impact of trade preferences on multilateral tariff cuts: Evidence for japan. *Journal of the Japanese and International Economies*, 38:31–51, 2015.
- D. J. Kuenzel, R. R. Sharma, et al. *Preferential Trade Agreements and MFN Tariffs: Global Evidence*. Wesleyan University, Department of Economics, Public Affairs Center, 2019.

- N. Limão. Preferential trade agreements as stumbling blocks for multilateral trade liberalization: Evidence for the united states. *American Economic Review*, 96(3):896–914, 2006.
- D. S. Lucas and C. J. Boudreaux. National regulation, state-level policy, and local job creation in the united states: A multilevel perspective. *Research Policy*, 49(4):103952, 2020. ISSN 0048-7333. doi: <https://doi.org/10.1016/j.respol.2020.103952>. URL <https://www.sciencedirect.com/science/article/pii/S0048733320300329>.
- J. Mai and A. Stoyanov. The effect of the canada-us free trade agreement on canadian multilateral trade liberalization. *Canadian Journal of Economics/Revue canadienne d'économique*, 48(3):1067–1098, 2015.
- M. S. Manger and K. C. Shadlen. Political trade dependence and north–south trade agreements. *International studies quarterly*, 58(1):79–91, 2014.
- E. Ornelas. Rent destruction and the political viability of free trade agreements. *The Quarterly Journal of Economics*, 120(4):1475–1506, 2005.
- E. Ornelas and M. Ritel. The not-so-generalised effects of the generalized system of preferences. *The World Economy*, 43(7):1809–1840, 2020.
- Ç. Özden and E. Reinhardt. The perversity of preferences: Gsp and developing country trade policies, 1976–2000. *Journal of Development Economics*, 78(1):1–21, 2005.
- A. Postigo. Explaining utilization of free trade agreements by sectoral interests and binding of unilateral concessions. *International Development*, 148:1–63, 2013.
- A. K. Rose. Do we really know that the wto increases trade? *American Economic Review*, 94(1):98–114, March 2004. doi: 10.1257/000282804322970724. URL <https://www.aeaweb.org/articles?id=10.1257/000282804322970724>.
- A. Subramanian and S.-J. Wei. The wto promotes trade, strongly but unevenly. *Journal of international Economics*, 72(1):151–175, 2007.

M. Tomz, J. L. Goldstein, and D. Rivers. Do we really know that the wto increases trade?
comment. *American Economic Review*, 97(5):2005–2018, 2007.