Tariffs, Retailers, and Consumer Prices

Kyung In Hwang*

(JOB MARKET PAPER)

November 4, 2018

Abstract

This paper examines how retail market powers influence the impact of tariff cuts on consumer prices in the context of Korea's tariff reduction since the mid-1980s. Using consumption-item-level consumer price index (CPI) and product-level tariff data with a new matching strategy, I empirically find that tariff cuts incompletely pass through to consumer prices. Motivated by a simple theoretical framework based on an oligopolistic retail competition model with a nested CES demand, this paper considers markups charged by retailers as a new channel by which consumer prices do not fully react to tariff cuts. To identify the retail markup channel, I firstly find that tariff pass-through to consumer prices is deteriorated in regions with greater retail market concentrations. Furthermore, this paper measures market-specific retail markups proxied by average price-cost margins, and quantifies the degree to which retail markups exacerbate tariff pass-through using an instrumental variable strategy. The instrument variable estimation reveals that a 10 percentage point increase in retail markups contributes to 1.73-1.96 percent lower tariff pass-through. My results are robust with estimation based on four-year census data by which I account for store-level variations and geographical market heterogeneities in retail markup measure. I also estimate the retail markup channel augmented by spatial autoregressive (SAR) model to control for CPI's spatially cross-section dependence, and the results are robust.

*School of Economics, University of Nottingham. Contact: kyung.hwang@nottingham.ac.uk. I am highly indebted to Sourafel Girma and Zhihong Yu for invaluable guidance, support and encouragement. I am grateful to Giammario impullitti, Holger Görg, Markus Eberhardt, Alejandro Riano, Holger Breinlich, Beata Javorcik, Halis Yildiz, Julia Cajal Grossi, and seminar and conference participants at the University of Nottingham, European Trade Study Group 2017, the Fall 2017 Midwest International Trade Conference, the 11th FIW International Economics conference for helpful questions and comments. All errors are my own.

1 Introduction

Standard international trade theories typically assume that producers sell goods directly to final consumer markets. In the real world, however, a very large portion of the seller-buyer transactions are conducted through retail distribution networks. The retail intermediation is, thus, likely to act as a trade friction and would substantially affect welfare gains of trade. Despite its making practical sense, international economics literature has little paid attention to the retail sector. This paper aims to explore the interconnection between international trade and the retail sector, focusing on examining how retail market powers influence the impact of trade liberalisation on consumer prices.

Domestic price level declines when a country implements trade liberalisation. It is verified by the fact that the removal of trade barriers leads to decrease in prices of foreign products and thus results in fall in domestic prices. Due to the clarity of the relationship, it has been little studied to quantify to what extent tariff reduction affects consumer prices. Yet, it is essential to shed a new light on the question of tariff pass-through into consumer prices. Tariff pass-through directly relates to macro perspective topics which have long been central questions in international trade. How large are welfare gains from trade liberalisation? How does trade reform affect inequality and poverty? The answers for these questions crucially depend on how trade liberalisation influences consumer gains through the price channel. The price channel of trade liberalisation is also a critical issue from a policy perspective. Although firms have been at the center of attention from trade economists, consumer welfare is also affected by trade liberalisations, as much as the degree to which consumer prices respond to reductions in tariffs.

It is widely regarded that sensitivity of consumer prices is proportional to changes in tariff rates. However, it is only feasible if there is no market friction from borders to consumer markets, which has been frequently rejected by a recent body of international economics literature. In particular, previous literature does not account for the role of retail firms in determining tariff pass-through. Considering that retailers distribute products from manufacturers or borders to consumer markets, tariff pass-through is highly likely to be influenced by pricing strategies of these intermediaries. This paper investigates the role of retailers as a channel through which consumer prices incompletely respond to tariff cuts.

To estimate the extent to which tariff affects consumer prices, I consider tariff reductions that took place in the Republic of Korea since the mid-1980s. Korea is an idealized case for this study for two reasons. First, Korea has aggressively opened the economy by reducing its tariff barriers. Interestingly, both degrees and timing in tariff reductions are varied across sectors in Korea. In fact, low time frequency in tariff data (i.e., tariffs change at most once a year) is considered as one of barriers to identify the price effect of trade reforms, which has not, by contrast, been an issue in identifying pass-through of exchange rates varying very frequently. In this sense, heterogeneous degrees and timing across sectors shown in Korean tariff cuts help to overcome the obstacle. Second, in parallel with tariff reductions, Korea's retail sector opened to foreign firms in 1996. The retail FDI liberalisation has transformed the landscape of Korea's retail sector. Most importantly, Korea has experienced a massive reallocation of retail market shares from small retailers to large-scaled retailers, thus leading to changes in the retail market structure. For these reasons, I exploit the Korean case in this paper.

This paper uses Korean consumer price index (CPI) data. Korean CPI has one great merit that it provides not only sector-level but also region-level price information since 1985. I take advantage of the three-way variations (i.e., sector, region, and year) both to estimate tariff pass-through and to identify whether retailers play a role in imperfect sensitivity of consumer prices to tariffs. I also make use of tariff data gathered over the period of 1988 to 2012. Since tariffs are provided at the disaggregated product level (HS 6-digit), it is essential to deal with the issue of matching product-level tariff data with sector-level CPI data. I match two data sets by using a bunch of existing concordances across product classifications established by the United Nation Statistics Division(UNSD). Through the matching procedure, I link CPI to tariffs, not relying on my own arbitrary decisions but relying on qualified and precise UNSD's concordances.

To test the effect of tariffs on consumer prices, I run a regression with sector, province and year fixed effects, along with observable controls to capture time-varying confounders. I find that changes in tariffs are not completely reflected in changes in consumer prices: a 10 percentage point decline in tariff rates is associated with 0.60 percent decrease in CPI. Interestingly, my finding is consistent with previous empirical studies on tariff pass-through such as De Loecker et al. (2016) and Topalova (2010), but the magnitude is slightly smaller than those papers. The discrepancy seems to indirectly indicate the existence of retail channel through which tariff pass-through is exacerbated between importers and consumer markets. It is because that CPI is the far downstream price while wholesale prices explored by Topalova (2010) and manufacturing factory gate prices investigated by De Loecker et al. (2016) are upstream relative to consumer prices.

As commonly concerned by the literature, tariffs may be determined by unobserved political economics factors, which would in turn lead to the endogeneity issue. That is, sectors with high performances would be able to lobby aggressively to the government, and the sector would have favourable tariff rates. If considering that the early stage of Korea's tariff reductions was driven by governmental policy decisions, endogeneity embodied in tariffs should be seriously taken into account for in the Korean context. I assess bias from endogeneity in tariffs by conducting a sensitivity analysis proposed by Oster (2017). I find from the sensitivity analysis that the entire effect of tariffs on CPI is unlikely to be driven by the bias from political economics factors. Thus, sensitivity analysis confirms that, while the precise estimate is still unknown, there should exist a causal relationship between tariffs and consumer prices.

Why are tariff reductions imperfectly implemented in consumer prices? Why do tariff cuts pass through to consumer prices less than what is expected? From a theoretical point of view, the answer is less clear. In this paper, I construct a simple theoretical model to explain why retailers negatively affect tariff pass through. Tariff reductions would not completely pass through to consumer prices if retail firms respond to the marginal cost shocks in a manner that they increase their markups while decreasing retail prices modestly. To uncover the mechanism in a theoretical framework, I consider an oligopolistic retail competition model with two-level CES demand system based on Dornbusch (1987) and Atkeson and Burstein (2008). The theoretical framework allows for *variable markups* across retailers, and shows that retail firms with higher market share is related to lower demand elasticity and thus higher markups. Relying on the setting, I demonstrate that retail markups lower tariff pass-through.

Being motivated by the theoretical framework, I empirically investigate the relationship between retail market concentrations and tariff pass-through. Since retail firms, on average, charge higher markups in highly concentrated markets, it is worth testing whether the effect of tariffs on CPI is deteriorated in markets where the retail sector is highly concentrated. To test it, I construct the Herfindahl index for the sector which varies across provinces and years. Then, I augment the Herfindahl index and its interaction with tariffs to the tariff pass-through regression specification explained earlier. As emphasised before, the empirical challenge is that tariffs are endogenous. To deal with the issue, I focus on the change within sector-year clusters by including sector-year fixed effects in the regression. Put differently, sector-year fixed effects are able to eliminate any kinds of sector-year specific confounding effects certainly including political economics components. I find that a 10 percentage point higher Herfindahl index causes 0.18 percent lower tariff pass-through rates.

While I show that retail market structure has a significant impact on CPI's sensitivity to tariffs, the question of whether markups charged by retailers affect tariff pass-through has not solved yet. To reveal correctly the role of retail channel, the focus should be on the impact of *retail markups* on tariff pass-through. However, it is not straightforward to measure province-level retail markups due to lack of available data. While recent papers proposed a methodology to construct econometrically estimated markups (e.g., De Loecker and Warzynski (2012)), it is extremely data demanding that it is hard to apply it to this study. Instead, my markup measure is proxied by price-cost margins. To construct them, I use KOSTAT(Statistics Korea; Korea's office of national statistics)'s data on the retail sector from which I get access to province-level aggregate data necessary for measuring price-cost margins.

Previously, I only care about tariff's endogeneity, whereas I do not for the Herfindahl index because market structure is assumed to be given exogenously as is done in the literature. On the other hand, when retail markups are adopted, they should be regarded as an endogenous variable as retail markups are not randomly determined in the regression of CPI on tariffs. Specifically, retail firms are more likely to open stores in regions where CPI is higher, thus leading to higher competitions and lower retail markups in those regions. To address the endogeneity, I propose two new instrument variable strategies. The first instrument variable is the province-level Herfindahl index for the *wholesale* sector. The instrument strategy crucially relies on the vertical relationship between retailers and wholesalers. While I focus on a horizontal market structure as a retail channel of imperfect pass-through, retail firms' vertical relationship with wholesalers certainly affect their pricing behaviours. In this sense, I use the wholesale Herfindahl index to instrument retail markups. Moreover, it is clear that wholesale market structures influence consumer prices only via retailer's pricing strategies because wholesalers are disconnected to consumer markets. The second instrument variable for retail markups is ten-year lagged populations. The instrument is based on the insight that market sizes are one of crucial determinants of retailers' market entry and location decisions. While populations are sometimes considered as being exogenously given in the economics model, I make use of very long lagged values because they make the variable far more uncorrelated to current consumer prices. From the IV estimation, I find strong evidence that retail markups have negative impacts on tariff pass-through: a 10 percentage point increase in intermediary markups causes 1.73-1.96 percent lower tariff pass-through.

Empirical findings so far may rest on how reliable retail markup measures (i.e., price-cost margins) are. Two issues may be raised. First, my price-cost margins are measured by a simple ratio of aggregate profits to aggregate sales at the province level. That is, there is no consideration of store-level variations. Second, geographical heterogeneities within a province are not accounted for, even though provinces are too spacious to be a single geographical market for a retailer. To overcome the issues, I investigate the retail markup mechanism focusing on KOSTAT's census data for the years of 1996, 2001, 2005 and 2010. Because, for the four years, store-level information is accessible, I first construct weighted average price-cost margins of within-province stores using census data. Census data provides detailed geographical identifiers such as county and town codes. Using county codes, I further build weighted mean markups with county-level market heterogeneities controlled by the fixed effect estimation approach. Limiting to four census years and exploiting two new price-cost margins, I confirm that retail markups truly affect tariff pass-through negatively. Lastly, I incorporate spatial autoregressive (SAR) model to retail markup channel to control for CPI's cross-section dependence at the region level, and I find that the results are very robust.

This paper is related to empirical studies on trade and prices. Porto (2006) investigates price effects of tariff reforms using consumer expenditure survey data in Argentina. Broda and Romalis (2008) study whether imports from China reduce cost of living in the U.S. using home-scanner data. Faber (2014) finds cheaper access to the U.S. inputs reduces relative prices of high quality products in Mexico. My study differs from them in that I relate CPI which is comprehensive of all industries to product-level (HS 6-digit) tariffs, and, to my knowledge, this paper is the first study to match CPI with disaggregated tariff data.

This paper also complements literature by providing evidence that changes in tariffs are not perfectly reflected in prices. Feenstra (1989) shows incomplete tariff pass-through happens if foreign exporters are under imperfect competitions. Nicita (2009) studies the impact of tariff reductions on local prices and shows the price effect differs across regions. Atkin and Donaldson (2015) also present that geographical distance to borders matters to tariff pass-through. Han et al. (2016) show the size of private sector across regions influences the sensitivity of the price effects of tariff reductions in China. De Loecker et al. (2016) find that manufacturing firms' factory gate prices incompletely react to tariff reduction, because they absorb decline in marginal costs as their markups. Apart from previous literature, I newly suggest that retail markups are a crucial channel of imperfect tariff pass-through into consumer prices.

From a perspective of the retail sector, research on exchange rate pass-through are related to my work. For instance, Burstein et al. (2003) and Goldberg and Campa (2010) show that incorporating the retail sector counterfactually improves the predicted pass-through rates. The exchange rate pass-through literature has paid attentions to two channels related to retail sector. First, local distribution costs may play an important role in exchange rate pass-through. (e.g., Burstein et al. (2003), Corsetti and Dedola (2005)) Second, markup adjustments of retailers exacerbate exchange rate pass-through. (Nakamura and Zerom, 2010) Even in the exchange rate pass-through, while there is ample evidence that low exchange rate pass-through can be explained by local distribution costs, there is less support for retail markups. (Burstein and Gopinath, 2014) This paper contributes to the literature by adding new empirical evidence that retail markups matter for marginal cost pass-through.

The rest of the paper proceeds as follows. Section 3 describes policy background and data used in this paper. Section 4 empirically uncovers the causal relationship between tariff reductions and consumer prices. Section 5 provides evidence of the impact of retail markups on tariff pass-through. Section 6 shows results of robustness checks. Section 7 concludes.

2 Background

2.1 Trade Policy Background

It was not until 1984 that Korea turned to trade reform. The trade reform aimed to achieve two objectives: uniform tariff rates across sectors and considerable reductions in tariffs. Before 1984, Korea's tariff scheme had been very differential across sectors and had maintained high rates. Accordingly, Korea's economy had been dependent on non-traded sectors. The tariff reform, targeting at accomplishing two purposes mentioned above, improved resource allocations and contributed to the growth of traded manufacturing sectors. (Cheung and Hongsik, 2010)

Korea's tariff reform has been carried out in two phases. Each phase has its own end and feature. The first phase of Korea's tariff reduction was conducted from 1984 to 1994. It was a government-driven tariff reduction and focused on manufacturing sectors. Since tariff rates had been substantially differential across sectors, the tariff reform in the first period was implemented in a manner that the Korean government sets pre-determined target rates which are uniform across sectors. As a result of the first reform, Korea's average tariff rate of manufacturing products declined from 40% to 14%.[1]

In 1995, Korea joined the World Trade Organization (WTO), so that Korea had to reduce importing tariffs in order to keep WTO requirements. The second phase of tariff reforms was driven by the Korea's joining WTO. Interestingly, tariff cuts were mostly implemented in agricultural products in the second phase because manufacturing products were already low enough to satisfy WTO's requirements. The average tariff rate of agricultural products dropped by 16.7 percentage points from 1996 to 2010. Even since 1996, manufacturing sector, though, has experienced additional tariff reductions. Unlike the pre-1995 trade reforms by which uniform target rates were applied to the whole manufacturing industries, tariff reduction in the second phase focused on selective manufacturing industries. For instance, in 1997, tariff rates in 182 raw materials and intermediary inputs, which had accounted for a large share in total Korea's imports, declined. In 2000, the Korean government also reduced 43 IT (Information technology) products, because Korea had to comply with agreements by

[1] The first phase of Korea's tariff reforms can be further divided into two waves. The first wave in this phase was conducted from 1984 to 1988. Tariffs had been reduced over the period, targeting at 20% uniformly across final products. The second wave took place from 1989 to 1994. Target rates were far decreased in this period to 8% across final products. (Cheung and Hongsik, 2010)

ITA (Information Technology Agreement).

In short, Korea's tariff cuts took place over the long period, and can be divided into two phases, each of which has a respective purpose. As a matter of fact, it is unlikely that tariff reductions took place exogenously, suggesting that endogeneity issue may be addressed when estimating the impact of tariffs on consumer prices. I will elaborate on how I deal with endogeneity in tariffs in section 4 and 5.

2.2 Korea's retail sector

In Korea, the retail sector offers a substantial contribution to the national economy. As Figure 1 shows, the sector is the second largest sector in Korea's economy. The share of the retail sector in Korea's total GDP is 8.2% in 2013. While it is about one third of manufacturing sector(28.4%), GDP share of the sector exceeds both the construction(4.5%) and the agricultural(2.1%) sectors in Korea. With regard to employments, the retail sector is far more important in Korea's economy. Figure 1 also presents that the retail sector accounts for 14.6% of Korea's total employments. Although it is fewer than the manufacturing employments, the differential in employment shares between the manufacturing and the retail sector is only 2.1% points. Still, the retail sector hires even more people than the construction sector(7.0%) as well as the agriculture sector(6.1%).

It was not until 1996 that Korea's retail sector turned to being a part of the global market. Foreign direct investment in Korea's retail market was completely liberalised in 1996, and multinational retail firms have entered Korea. As a matter of fact, the retail sector prior to opening the market had been one of the least developed sectors in Korea. Before 1996, mom-and-pop stores and small local supermarkets accounted for a majority of Korean retail sale (Cho et al., 2015). According to the KOSTAT (Korea's national office of statistics), the average number of workers per retail store, which indexes the size of retail stores, is 1.9 - 2.1 before 1996, and it is remarkably lower than other countries.[2] After 1996, however, the Korean retail sector began to experience dramatic changes. There are three noticeable structural changes in the Korean retail sector.

The first structural change appeared in Korea's retail sector is that total sectoral sales become almost doubled from 1996 to 2010, and that the expansion was mainly driven by large-scaled retail firms. Figure 2 illustrates the evolutions of total retail sales from 1996 to 2010. As shown by panel (a), total retail sales were 141.6 trillion Korean won (KRW) in 1996 but it increases to 262.8 trillion KRW in 2010.[3] Panel (a), however, demonstrates that time

[2]For instance, the average worker per retail establishments in Japan already reached to 3.9 in 1985 (Kim and Kim, 2003).

[3]The sales amounts in Panel (a) of Figure 2 are all expressed in real values deflated by Korea's

trends in total sales differ between corporation-type retail firms and sole traders(i.e., self- and family-owned retailers). Specifically, while corporation-type retailers show an almost identical trend with total sectoral sales, sole traders representing small-scaled retail firms have a decreasing trend during the period of 1996 to 2005. In terms of market shares, retail corporations show an increasing trend as depicted by Panel (b), whereas sole traders move in the opposite direction. In particular, market shares of retail corporations began to surpass those of sole traders from 2001.

Second, although total sales have expanded since 1996, the number of retailers has shown a dramatic plunge over the period. In Figure 3, I respectively plot total numbers of entire retailers (Panel (a)), corporation-type retailers (Panel (b)) and sole traders (Panel (c)) across years. Panel (a) of Figure 3 presents that the number of retail firms was 768,875 in 1996 whereas it falls to 579,356 in 2010. The massive reduction is driven by sole traders as illustrated in Panel (c). On the other hand, the number of corporation-type retailers even soars from 13,828 in 1996 to 32,974 in 2010.

The third salient feature is the increasing prevalence of large discount stores. Large discount stores are a retail store format firstly introduced in the mid of 1990s in Korea. A typical store belonging the format has an area of 40,000 square metres of floor spaces, sells 50,000-100,000 product items and hires 100-150 employees (Cho et al., 2015). Large discount stores are analogous to hypermarket stores or superstores that sells general merchandises at low prices. Since multinational retailers have mainly operated stores of this format in the global market, abolishing the limits on foreign retail entry in 1996 contributed to the proliferation of the stores established by multinationals. At the same time, being forced to face a new circumstance, Korean conglomerates (so-called 'chaebol') that have run their businesses in the consumer-goods sector or department stores actively introduced and increased the new store format. As a result, there were only 17 large discount stores in 1996, but the total number of large discount stores increases to 420 in 2016.

The aforementioned three features, which have characterised the Korean retail sector since 1996, are indicative of economic consequences of retail globalisation in two folds. First, retail FDI may induce increase in competition in host-country retail markets. Figure 2 suggests that Korea's FDI liberalisation contribute to reallocating market shares from small retailers towards large-scaled retailers. Figure 3 also shows that there has been a large degree of firm churning in the retail sector since 1996. The structural changes indicates that fiercer competition induced by retail FDI liberalisation transformed the landscape of the retail sector.

Second, retail globalisation is associated with changes in market structure. I have noted

CPI(consumer price index).

that while total retail sales have considerably increased, the number of firms has highly decreased since 1996. Therefore, it is easily predicted that the entry of foreign retailers contributes to changes in retail market structure, a transition towards increase in market concentrations. Furthermore, if considering that large discount stores profoundly affect local retail markets as extensively shown by the so-called 'Wal-Mart effect' literature, I would expect that degrees of concentrations may differently be determined across local markets, depending on whether or not foreign stores are present in local markets.

3 Data

3.1 Price and Tariff Data

I make use of Korea's consumer price index (CPI) collected by the KOSTAT (Statistics Korea). The KOSTAT is a governmental agency which is responsible for national statistics. Korean CPI data provides not only sector-level but also province-level price information. More specifically, Korean CPI contains price indices across 72 sectors and 16 provinces since 1985.[4]

I use Korean CPI data for three reasons. Firstly, CPI is directly linked to this paper's research question. Previous literature focuses mainly on export or import unit prices to investigate tariff pass-through. In this study, however, I aim to evaluate the causal effect of tariff cuts on prices that consumers face. Considering that CPI is constructed by surveying sales prices at retail stores, it is highly suitable for this end. Secondly, regional variations in Korea's CPI play an essential role in this paper. The variation across regions allows us to investigate the role of retail intermediaries in tariff pass-through, because retail market structure measures used in this paper are established at the region level. Thirdly, Korea's CPI has an advantage that it provides substantial amount of sectoral price information as well. The sector-level variation is also crucial for this study in the sense that tariff data varies across sectors. As far as I know, there are few countries that publicly release such a substantial amount of sector-level CPI information.

Korean CPI data is not a balanced panel. It is because the number of provinces surveyed for CPI has incremented over years. It started with 6 provinces at first, rose to 15 in 1990,

[4] In fact, *sectors* may not be an appropriate terminology for CPI categories, in the sense that CPI categories are much closer to consumption items that are components of the consumption basket, rather than physical products like Harmonised System (HS) codes. Nevertheless, I use the terminology, sectors, for CPI categories for the convenience reason. Above all, as I will explain in section 3.2, I successfully match each CPI category to HS codes with a novel matching strategy that I devise. In this regard, I think there is no problem in calling CPI consumption categories *sectors*. Throughout this paper, I use the terminology CPI sectors instead of CPI consumption-item.

and became 16 from 1995 up to the present. Although the KOSTAT began with only 6 provinces for the first 5 years(1985-1989), the 6 provinces actually accounted for 43.2% of Korea's total GDP.[5] Hence, it seems that even CPI during initial five years is sufficient enough to represent the Korean economy.

I also utilize Korea's tariff data set drawn from UNCTAD-TRAINS (United Nations Conference on Trade and Development-Trade Analysis and Information System). It contains not only MFN applied tariff rates but also import flows since 1988, both of which are based on six-digit HS (Harmonized System) product level. Unfortunately, it provides Korean tariffs with missing values in nine years (1991, 1993-94, 1997-98, 2001, 2003, 2005, 2008). I fill the gap with another data source, WTO-IDB (World Trade Organization-Integrated Data Base) which also contains both MFN applied tariffs and imports from 1996 to 2012. I confirmed that two sources have identical information during the overlapped years by both data sets. So, I combine both data sources and use it as one dataset. Still, even using two sources of data does not recover three-year tariffs (1991, 1993 and 1994). In short, tariff data used for this analysis is applied MFN tariffs across all products at the HS six-digit level and it ranges from 1988 to 2012 with three-year missing values.

Among a total of 72 sectors classified by Korean CPI, this paper restricts to 26 traded manufacturing sectors which are able to relate to tariffs. In CPI sector classifications, 30 sectors are classified as service sectors(e.g., housing services, insurances, restaurants and hotels), so I drop them because they have nothing to do with tariffs. I further exclude three manufacturing sectors - electricity, gas, and newspapers and periodicals, since there are few records in which products in those sectors have been imported from foreign countries in my dataset. I consider the three sectors as non-traded ones.

I also exclude 13 traded agricultural sectors for two reasons. Firstly, it is clear that pre-1996 tariffs of agricultural products are inappropriate for an indicator of protection. According to Table 1, there appears a sudden and abnormal surge in agricultural tariffs between 1995 and 1996. As a matter of fact, it results from the so-called tariffication. Tariffication was one of UR (Uruguay Round)'s agreements, which required developing countries to convert agricultural non-tariff barriers(NTBs) into tariffs. The high increase in tariffs in 1996, therefore, has nothing to do with increase in protection level. Rather, it is an outcome of the action intended to transform NTBs into quantifiable measures in order for Korea to render its agricultural trade policy more transparent. Secondly, in Korea, special safeguard mechanism (SSM) has often been carried out in agricultural sectors over years. SSM is a tool that

[5] The share of 6 provinces in Korea's total GDP is calculated on the basis of 2003. I compute the proportion from KOSTAT's regional GDP data. Since the KOSTAT releases province-level GDP from 2003, I inevitably use the share of 6 provinces as of 2003.

allows developing countries to raise tariffs temporarily to deal with import surges or price falls. It explains why there are some agricultural tariff observations that have sky-scraping values in my dataset. Therefore, I do not include agricultural sectors in this paper.

As noted above, it is not available to get access to tariff data in 1991, 1993 and 1994. Recall that a large portion of decrease in manufacturing tariffs took place before 1996. Thus, some may argue that pre-1996 tariffs contain meaningful information and that missing in those data would have a harmful influence on regression results. For the reason, I fill in the missing observations by a linear interpolation. The rationale for using linear interpolation is that the missing seems to be "Missing at Random (MAR)" case, which means it is unlikely to have specific intentions by data providers. More importantly, considering that before 1996, Korea carried out tariff reductions in a way that it just followed a *pre-determined* reduction plan which has a linearly monotonic decreasing trend, I expect that the linear interpolation would be able to revive missing data as correctly as possible. Table 2 shows the target tariff rates set by the Korean government from 1988 to 1994.

3.2 Matching CPI with Tariffs

My empirical analysis critically rests on matching CPI data with tariffs. Matching of both data sets is significantly demanding work as two sources of data have different levels of aggregation. In particular, CPI is highly aggregated (i.e., CPI has totally 72 sector-level categorises), whereas tariffs are very disaggregated (i.e., Tariff data has individual values across HS 6 digit level products). More seriously, each data follows different classification standards. Trade data exists on the basis of HS codes, whereas CPI builds on a different consumption-item classification standard, called COICOP (Classification of Individual Consumption According to Purpose). It would seem that such matching difficulties lead trade economists to rarely estimating consumer price effects of trade reforms.[6]

In this paper, I construct a concordance between CPI and tariffs, which has, to my knowledge, not been generated before. The bottom line of my matching procedure is that I exploit existing United Nations Statistics Division (UNSD)'s concordances across related product classification standards. In fact, the UNSD has established and maintained various kinds of product classification standards, and it offers concordances between them as well. Unfortunately, the UNSD has not generated a concordance between CPI and HS codes. In this paper, I use UNSD's various concordances as a kind of *bridges* to connect CPI to tariffs.

Figure 1 illustrates the matching procedure that I exploit to relate CPI to tariffs. The

[6] Broda and Romalis (2008) is an exception. They construct a concordance between barcode-based product level prices and HS codes. However, their approach is different from this paper in the sense that they rely on the arbitrary matching based on authors' subject judgement and product descriptions.

procedure consists of four steps. The first step is to match CPI to COICOP. According to the KOSTAT, classification of consumer price index is fundamentally based on COICOP, which in priciple allows for one-to-one matching between CPI and COICOP.[7] Secondly, I relate the matched COICOP with CPC using UNSD's concordance table. In this step, all the matching is automatically conducted only relying on UNSD's concordance table. The next step is to match CPC with HS codes utilising related UNSD's concordance table. As is well-known, HS system has changed its verstion four times so far. For this reason, I firstly match CPC products with HS nomenclature 1996 version (HS96) because CPC is able to match only with HS96 through UNSD's concordance. Then, I match the results to other HS versions, using UNSD's concordances across HS nomenclature versions.[8]

The final step is to construct weighted average tariff rates across CPI sectors. Completing the matching procedure up to the third step, each CPI sector becomes having a great deal of corresponding product (HS six-digit) level tariff rates. Thus, I calculate the average of tariff rates across HS six-digit level products within each CPI sector. The weights that I utilize are product-level(HS 6-digit level) share of one-year lagged imports in total sectoral imports one year before. Although many of related papers has used current imports as weights for average tariff measure, my preferred weights for this study are imports one-year before. The reason is that determination of tariff rates may be affected by foreign imports, which potentially raises an endogeneity issue. Consequently, the formal specification of weighted average tariff of sector i at year t is

$$\tau_{it} = \sum_{j \in I_i} (weight_{ijt} \times \tau_{ijt})$$

where j indexes a 6-digit HS level product, I_i is a set of product j matched with sector i by my concordance table. As described above, the weight of product j and sector i at year t is defined as

$$weight_{ijt} = \frac{Imports_{ijt-1}}{\sum_{j \in I_i} Import_{ijt-1}}$$

[7]There are, though, four CPI sectors that cannot be directly matched to COICOP. The sectors all belong to the textile industry. The matching is difficult because I need to do match four CPI sectors(*Men's Clothes, Women's Clothes, Casual Clothes* and *Children Clothes*) to two COICOP categories(*clothing material* and *garments*). In this case, I link CPI to CPC (Central Product Classification), without using COICOP as a bridge. Note that, as illustrated in Figure 1, CPC is used as a bridge to connect COICOP to HS codes, which I explain in detail in the main text. Using UNSD's concordance between COICOP and CPC, I find CPC products corresponding to the two COICOP items. Then, I allocate each of matched CPC products to the four CPI sectors by examining descriptions of both CPI and CPC in detail.

[8]COICOP is a reference classification particular for consumption items. CPC is a product classfication based on physical characteristics of products. All the concordances developed by UNSD can be found in UNSD classification registry (http://unstats.un.org/unsd/cr/registry).

3.3 Data Summary

Table 3 shows the summary statistics of variables in the dataset. The dataset used in this study contains a total of 10,400 observations (26 sectors $\times$ 16 provinces $\times$ 25 years = 10,400). Table 3 presents that there are 3,770 missing values in CPI variable as Korean CPI initially started with only 6 provinces and the KOSTAT has gradually increased the number of provinces surveyed as noted above. The missing values in weighted average tariff data is due to lack of import data in 1987, one year before the sample period (1988-2002). Weighted average tariff rates in 1988 are unable to be established as I use one-year lagged imports as weights. Figure 5 illustrates a time trend in the logarithm of aggregate CPI, with a base year 2010 (CPI_{2010}=100). It shows that log CPI increases at the beginning (1988-1998), sluggishly rising or even flat in the middle (1999-2007), and rises again in the late (2008-2012) of sample period.

In Table 4, I report sector-level log consumer price indices over years.[9] The table illustrates that most sectors show increasing trends, while three sectors are exceptionally decreasing.[10] Interestingly, the three sectors are all classified into information technology (IT) industry. It seems that decrease in CPI of the three sectors is due to rapid IT technology innovations. In addition, the reduction of log CPI in three sectors may be a cause of the flat trend of aggregate (log) CPI between 1999 and 2007. In other words, the CPI decrease in IT sectors is so profound that it substantially influences the rest of sectors.

Table 5 describes provincial log consumer price indices over years.[11] Unlike Table 4, every province has experienced on-going increase in log CPI over the sample period. A marked feature appeared in Table 5 is that differences between provincial indices are minor. This feature contrasts with the high-level of differences between sector-level indices shown in Table 4. It clearly suggests that sector-level heterogeneity needs to be more seriously taken into account when regressing log CPI on tariffs. To the end, I use a sector fixed effect in the regression, and I also exploit sector-year specific control variables to account for the heterogeneity issue. I explain details of the fixed effect as well as controls in section 4.

Figure 6 presents time trends in average tariff rates over the sample period. I draw time trends for weighted average tariffs and simple average tariffs, separately. As well illustrated in Figure 6, two average tariff measures have almost identical trends over the sample period.

[9] A sector-level log CPI is calculated by averaging out all provincial log CPI within each sector. A sector-level standard deviation of log CPI, thus, measures the degree to which province-level indices are dispersed within each sector.

[10] Three sectors where log CPI has a downward time trend are *equipment for sounds and pictures*, *information processing equipment*, and *telephone and telefax equipment*.

[11] Each province-level index is the average of sectoral log consumer price indices within the province. Standard deviations present degree of variations in log CPI at the province level.

While average tariff rates in general show steadily downward trends, a slight increase appears in 1998. The abnormal rise in 1998 resulted from a temporary trade policy which reacted to Korea's financial crisis in 1997. Afterwards, tariffs has again been dropped up to 2012. As discussed in section 2.1, tariff reductions significantly took place before 1995, the year when Korea joined WTO member countries. Although it is still downward trend, tariff reductions after 1995 look very modest. The pattern also accords well with what I explain earlier on Korea's trade reforms. In section 2.1, I mentioned that the Korean government selectively reduced tariff rates for some of sectors. In terms of aggregate rates, the degree of reductions does not seem to be substantial. Yet, the selective changes in sector-level tariffs are reflected in the increase in deviations across sectors. Figure 4 illustrates tariff dispersions across sectors. Figure 4 shows that, while sectoral dispersions continuously decline up to 1995, it has been enlarged since 1996 because of the differential tariff policy across sectors.

4 Tariffs and Consumer Prices

4.1 Regression Specification

Let i index sectors, r denote provinces and t imply years. Equation (1) describes my preferred regression specification to estimate the degree to which importing tariffs have an influence on consumer prices in domestic markets.

$$\ln p_{irt} = \beta\tau_{it} + {Z_{it}}'\gamma + \delta_i + \eta_r + \lambda_t + \varepsilon_{irt} \tag{1}$$

where lnp_{irt} is the logarithm of the consumer price index (CPI) of sector i in province r and year t. τ_{it} is the tariff rate of sector i in year t, which is the weighted average across products within the sector as described in section 3.2. β is a coefficient of interest. It captures tariff pass-through rate which implies how much percentage of CPI would decrease (increase) when tariff rate declines (rises) by 1 percentage point. δ_i is sector fixed effects to account for unobserved time-invariant characteristics across sectors. The province fixed effect η_r is to capture differences across provinces such as geographical heterogeneities and labour costs. I control for macroeconomic fluctuations through λ_t.

In section 3.3, I emphasised the necessity of controlling for sectoral heterogeneity by showing Table 4. While I include sector-specific fixed effects (δ_i) in the regression equation above, it might be concerned that there are still time-varying unobserved sectoral heterogeneities. Obviously, it is impossible to include the sector-by-year fixed effects, because the fixed effects will make the variations in tariff rates disappeared. To account for the unobservables, I use three sector-year specific control variables (Z_{it}).

First, I employ import penetration ratio. Import penetration ratio, which measures the level of import competition, is defined as the ratio of imports to imports plus domestic productions in a given sector and a given year. Recall that a sector, in this paper, is defined as a consumption item which is basically a component of consumption baskets used in CPI construction and that I call it CPI sector. This implies that, to measure sector-level import penetrations, both imports and domestic productions are required to be calculated on the basis of CPI sectors. For the reason, I assign product-level imports to CPI sectors through my concordance table between HS codes and CPI, and then I compute sector-level imports by summing up all product-level imports belonging to the corresponding CPI sector. However, with regard to domestic productions, it is impossible to measure them across CPI sector-level due to lack of available data. Thus, I use consumption data as a proxy for the denominator of import penetration ratios, i.e., the sum of imports and domestic productions.[12] Although the proxy has a limitation that exports are disregared when I replace imports plus domestic productions by consumptions[13], I view consumptions as a proxy as good as I can, especially under lack of data in Korea.

I use world prices as the second control variable. International trade theories typically suggest that, in small (or price-taking) countries, domestic prices of foreign goods are affected by world prices. In this sense, many related studies include world prices in the tariff pass-through regression.(e.g., Nicita (2009), Feenstra (1989), Han et al. (2016))[14] As it is difficult to measure world prices, I exploit the U.S. export prices for the proxy of world prices, which is widely used in the literature (Han et al., 2016). I obtain both US export values and export quantities from UN Comtrade dataset, calculate export unit values (i.e. the ratio of export value to export quantity), and use it as the proxy for world prices. Since the U.S. export data exists at the HS 6-digit level in UN Comtrade, I calculate an average world price of each CPI sector weighted by product-level export share within each sector, and use

[12]The sources of consumption data are Household Income Expenditure Survey(HIES) and population census, all of which are collected by the KOSTAT. Regarding the classification issue, it is very suitable for this study, because the KOSTAT has constructed data at the CPI sector level. As a matter of fact, the HIES data does not contain total value of sectoral consumption but *average* value of it, which is a little weird to us. However, population census provides the number of households over years. I construct CPI sector specific consumptions, multiplying sectoral average by the number of households.

[13]When considering the relationship between domestic production and consumption, it can be expressed, in a very simple term, that $consumption = domestic\ production - (exports + imports)$. Therefore, replacing domestic productions plus imports by consumption may be problematic in that exports are not taken into account.

[14]In order for world prices to capture the confounding effect in the regression, there should be a relationship between Korean tariffs and world prices. It is not straightforward to answer whether world prices are correlated to Korean tariffs. A plausible answer is that the global trend toward trade liberalisations, which certainly influences world prices, has affected Korean tariffs. For instance, WTO regime, which acted as a driving force of Korea's trade reforms, have enormously affected on global prices of traded goods.

it as sector-level world price. When relating product-level exports to CPI sectors, I use my concordance described before.

Third, I exploit imports from China to Korea. Since 2001 when China joined WTO, share of China's products in Korea's total imports has rapidly increased. Specifically, the share rose from 17.4% in 1997 to 41.5% in 2007. Like many countries in the world, it has been a common phenomenon that *made-in-China* products are displayed in every corner of Korean retail stores. Thus, it is highly likely that Korea's CPI is influenced by the growth of imports from China. Further, it is clear that Korea's tariff policy is associated with the rapid expansion of imports from China. China's exports to Korea would not have increased that much if it had not been for Korea's tariff reductions. Put together, imports from China to Korea would be a good control variable.

In terms of standard errors for estimates, I use clustered standard errors at the sector-by-year level, which aim to control for spatially serial correlations in the error term. When using a spatial panel data, estimates may be at risk of being contaminated by the spatial error correlation. In fact, Table 5, where I report province-level average (log) CPI over years, suggests that there may be a certain co-movement of Korean regional consumer prices over time. On the other hand, there seems to be less concern about sector-level correlation among prices as presented in Table 4. Hence, if residual autocorrelation exists, province-level serial correlations, arguably, should be more taken into account than sector-level serial correlations when regressing log CPI on tariffs. In this regard, I choose clustered standard errors at the sector-by-year level so as to capture within-cluster (i.e., province-level) auto correlations.

4.2 The Impact of Tariffs on Consumer Prices

Table 6 presents the estimation results of Equation (1). The parameter of interest in Equation (1) is β which implies the effect of tariffs on log CPI. I report the estimates for β in the first row of the table. The pooling OLS results without any controls presented in column (1) show that $\hat{\beta}$ is statistically insignificant. Also, applying either sector fixed effects or province fixed effects to the regression makes no difference with the pooling OLS as reported in columns (2)-(3). However, $\hat{\beta}$ turns into a positive and significant estimate when I include year fixed effects in addition to sector and province fixed effects (column (4)). It may indicate that it is important to control for variations in years because, as shown in section 3.3, CPI and tariffs have monotonic time trends. Column (5), where three sector-year specific controls are additionally included, is the result of my preferred regression specification. The estimate for β in column (5) is highly significant and positive.[15]

[15] As mentioned in the previous subsection, standard errors are computed by clustering at the sector-by-year level. As a robustness check, I use standard errors clustered at different levels and see if results are

In terms of magnitudes, column (5) shows that a 10 percentage point fall in tariff rates is associated with 0.60 percent decrease in consumer prices. Interestingly, using the context of India's trade reform from 1989 to 1997, De Loecker et al. (2016) presents a 10 percentage point fall in tariffs leads to a 1.29 percent reduction in factory gate prices of manufacturing firms. Topalova (2010) also shows that 10 percentage point decline in tariff rates in India results in fall in wholesale prices by 0.96 percent during that period. Two aspects are noteworthy. First, although I analyse a different country, Korea, in this study, my results do not look far from research about Indian trade reforms. Second, the size of the effect in this paper is slightly smaller than De Loecker et al. (2016) and Topalova (2010), but the discrepancy rather seems to accord to what this paper aims to investigate. In this paper, I evaluate tariff pass-through to CPI that is the downstream price paid by consumers at the retail level. On the other hand, De Loecker et al. (2016) exploit manufacturing factory gate prices and Topalova(2010) utilizes wholesale prices, both of which are upstream prices relative to CPI. Therefore, the smaller effect in this paper than those in De Loecker et al. (2016) and Topalova(2010) may indicate that there are retail market frictions by which tariff pass-through rates are deteriorated.

Even though the impact of tariffs on (log) consumer prices is estimated to be significant and positive as expected, it is hard to believe that $\hat{\beta}$ in column (5) is correctly estimated. It is because the estimation so far does not address an econometric challenge when tariff rates are used as a treatment variable: the endogeneity issue embodied in tariffs. In terms of this study, when estimating tariff pass-through, it is essential to control for endogenous determinations of tariff rates with respect to consumer prices. In particular, it is concerned from the trade liberalisation literature that tariff determinations are likely to be influenced by sector-level political powers. This seems to be quite relevant to the context of Korea's trade reforms. Korea's tariff cuts, especially from 1988 to 1995, were driven by governmental policy decision as explained in section 3. It is, thus, probable that tariffs are affected by sectoral political pressures or aggressive lobbies on policy decision makers. Moreover, the sector-level political powers highly depend on sectoral outcomes which affect sector-level consumer prices. In this regard, political powers across sectors can be a confounding factor in my regression model. While I use sector-year specific control variables(Z_{it}) plus sector, province and year fixed effects in Equation (1), the estimate $\hat{\beta}$ in column (5) of Table 6 would still be biased if the political confounding factor exists as an unobservable varying across sectors and years.

In this subsection, I focus on the sensitivity analysis built on insights from Oster (2017),

changed. I find that, whether clustering at the province-by-year level or clustering at the sector-by-province level, estimated standard errors for every coefficient in Equation (1) are almost identical to those in Table 6.

which assesses how large the bias due to unobservable confounders is. Oster (2017) suggests an intuitive way to examine the bias caused by unobservable confounding effects under so-called a proportional selection assumption (PSA), an assumption that selection on *unobservables* is proportional to selection on *observables*.[16]

There are two reasons for concentrating on the sensitivity analysis. First, it is extremely difficult to find proper econometric methodologies (e.g., instrumental variable strategy) to address the endogeneity, in particular in the context of Korea's tariff cuts. In section 3, I explained that Korea's tariff reductions have been carried out in two phases, each of which has its own purpose and feature. The dichotomy would make it hard to find econometric approaches suitable for the entire sample period. Second, while this analysis cannot give an answer to what degree tariffs affect consumer prices, it will certainly give an answer a question about whether causal effects of tariffs on CPI exists. Considering that the objectivity of this paper is not estimating tariff pass-through rates themselves, but identifying whether and how large they are affected by retail markups, it is more important to control for endogeneity in tariffs in the *retail channel regression* that I will specify later. Most importantly, I find a way in which I deal with the endogeneity issue when estimating the retail channel regression.

Table 7 shows the result of sensitivity analysis. Column (1) shows the regression results of Equation (1) from which I exclude observables (δ_i, η_r, λ_t, Z_{it}). Column (2) presents regression results with full sets of controls. Based on the formula defined by Oster (2017) (see footnote 14) and estimation results reported in columns (1)-(2), I calculate a maximum of the bias as -0.0208. The sensitivity analysis indicates that there is no evidence that CPI effects of tariffs is entirely driven by unobserved political powers. It is because the bias (-0.0208) only accounts for 34.7% of the point estimate for β (0.0599) in column (5) of Table 6.

[16]PSA is a key assumption in Oster's (2017) approach. The implication of PSA is that the selection on observables delivers considerable information about selection on unobservable confounding factors. Let W index observable explanatory variables and $W^{'}$ denote unobservable confounding factors. The *proportional selection assumption* (PSA) implies $\frac{Cov(X,W^{'})}{Var(W^{'})} = \delta\frac{Cov(X,W)}{Var(W)}$, where X is the explanatory variable of interest (i.e., tariff rates in this paper), and δ is a measure of how much correlated both selections are. Now, I consider regressions as below,

$$
\begin{aligned}
y &= \beta^F X + W + W^{'} + \varepsilon_F \\
y &= \beta^1 X + \varepsilon_1 \\
y &= \beta^2 X + W + \varepsilon_2
\end{aligned}
$$

and ${R^2}_{max}$ denotes a hypothetical R-squared of the first regression which fully includes not only observables (W) but also unobservables ($W^{'}$). ${R^2}_1$ is R-squared of the model without both observables and unobservables (the second regression), and ${R^2}_2$ is R-squared with obeservables (the third regression). In this set-up, Oster (2017) suggests that the unobservable confounding effect can be measured by

$$\text{Confounding Effect} = \delta\frac{(\beta^1 - \beta^2)({R^2}_{max} - {R^2}_2)}{({R^2}_2 - {R^2}_1)}$$

Nonetheless, the sensitivity analysis also suggests that political economics factor needs to be taken into account for further regressions, because the bias can be seen, in a different angle, as much as 34.7% of the total effect of tariffs on CPI. In the following section, I discuss how I address the endogeneity problem when I investigate retail channels by which tariffs incompletely pass through to consumer prices.

I experiment upon an alternative econometric specification used in several empirical papers on tariff pass-through. Initiated by Feenstra (1989), there are many studies (e.g., Nicita (2009), Marchand (2012), Han et al. (2016)) which exploit a regression specification to relate prices to tariffs, based on the following idea: in small countries, the relationship between tariffs and domestic prices can be basically expressed as $p_{it} = p^*_{it}(1 + \tau_{it})$, where p_{it} is domestic prices, p^*_{it} is world prices and τ_{it} is tariff rates. In sequence, if I take a logarithm on both sides of the equation, then it would generate $\ln p_{it} = \ln p^*_{it} + \ln(1 + \tau_{it})$. In this sense, I re-estimate Equation (1) with the tariff variable replaced by $\ln(1 + \tau_{it})$ as below

$$\ln p_{irt} = \beta \ln(1 + \tau_{it}) + {Z_{it}}' \gamma + \delta_i + \eta_r + \lambda_t + \varepsilon_{irt} \tag{2}$$

Table 8 shows that the replacement makes almost no difference with the original results in Table 6. Exceptionally, the impact of $\ln(1 + \tau_{it})$ on CPI is highly significant and positive, but the magnitude (0.212) becomes far larger than that of original model (0.0599). This result is quite obvious because alternative specification (2) is no other than a rescaling version of Equation (1) in terms of the tariff variable.

5 Retail Markup Channel

In the previous section, I find that tariff reductions lower consumer prices in Korea. Nonetheless, the effect does not seem to be sizeable that much. As mentioned before, I show that a 10 percentage point decline decline in tariffs is associated with 0.60 percent decline in consumer prices (see Table 6, column (5)). From 1989 to 2012, (weighted average) tariff rates fall, on average, by 7 percentage points. My estimate above indicates that the estimated average price decline by 0.42 percent (= 7 × 0.06), which is a very small effect of tariffs on consumer prices. It is a departure from what traditional trade literature considers tariff pass-through as being perfect, meaning that change in consumer prices is proportional to change in tariffs. Section 4, however, provides no answers of what the driving forces through which tariff reductions incompletely affect consumer prices are. This section investigates the role of retailers' market powers in determining the imperfect tariff pass-through into consumer prices.

5.1 A Simple Theoretical Framework

A traditional set-up in international trade models is that prices are determined by movement of demand and supply at the *manufacturer* level. Yet, a few papers have paid attention to behaviours of *retailers* in price determination. For instance, Nakamura (2008) finds that barcode-level price variations result mainly from retail- rather than manufacturing-level, suggesting that the retailers' pricing is a crucial factor in determining consumer prices. In this respect, it is worth studying whether tariff pass-through into consumer prices is affected by retail market powers. Intuitively, decline in tariffs would not fully pass through to consumers if retail firms with market powers respond to tariff reductions in a manner that they raise their markups to a significant degree while decreasing their retail prices modestly.

To guide my empirical analysis, I describe how tariff pass-through is deteriorated by retailers' market powers in the simplest partial equilibrium model based on Dornbusch (1987) and Atkeson and Burstein (2008), and an extention of Dornbusch (1987), Hong and Li (2017). Suppose a retailer sells a variety i in an oligopolistic market and it is assumed to take its marginal cost as given. The retailer, thus, sets the price of a variety i as a markup over marginal cost as the following:

$$p_i = \frac{\epsilon_i}{\epsilon_i - 1} w_i \tag{3}$$

where ϵ_i is the elasticity of demand of retailer i and w_i is wholesale cost paid by retailer i. I assume that wholesalers import foreign products, they provide only imported goods to retailers, and that they pay tariffs in order to buy foreign goods. Assuming, too, that the wholesale sector is perfectly competitive, they set wholesale price w_i, such that $w_i = c\tau$ where c is the unit price of imported product and τ is a tariff rate for the imported good.[17] Therefore, the retail price can further be expressed by

$$p_i = \frac{\epsilon_i}{\epsilon_i - 1} \cdot c \cdot \tau \tag{4}$$

Equation (4) indicates that while the process of supplying goods to consumers involves vertically multiple steps, which gives rise to so-called double marginalisations initially analysed by Spengler (1950), this model can avoid the issue by assuming the wholesale sector as being perfectly competitive. The objective of the setting is that I focus on the relationship between tariff pass-through and retailer's market powers. Rewriting Equation (4) as a full

[17] In fact, the more precise expression for wholesale price is $w_i = c(1 + t)$ where t is ad-valorem tariffs. For the convenience reason, I denote $\tau \equiv 1 + t$ and thus wholesale price can be defined as $w_i = c\tau$ as above, which never change the implications of the model. When I consider specific tariffs rather than ad-valorem tariffs, wholesale price w_i will be $c + \tau$. Though choosing whichever types of tariffs, results are not changed.

log differential,

$$d \ln p_i = d \ln \frac{\epsilon_i}{\epsilon_i - 1} + d \ln \tau \equiv d \ln \mu_i + d \ln \tau \tag{5}$$

where I assume that c is constant and μ_i denotes retail markups $(=\frac{\epsilon_i}{\epsilon_i-1})$.

Equation (5) reveals how retail markups relate to tariff pass-through. Suppose retail markets are perfectly competitive and μ_i is thus zero, then tariff pass-through will definitely be one $(\frac{d\ln p_i}{d\ln \tau} = 1)$, which I call a perfect pass-through case. Consider that retail markup is not variable with respect to tariffs $(\frac{d\ln \mu_i}{d\ln \tau} = 0)$. In this case, tariff pass-through will also be zero. Further, tariff pass-through will be more harmed if retailers charge more markups in response of tariff reductions $(\frac{d\ln \mu_i}{d\ln \tau} < 0)$.

To make it explicit that variable retail markups affect tariff pass-through from a theoretical perspective, I consider an oligopolistic competition model of variable markups with a nested CES demand system following Atkeson and Burstein (2008).[18] I assume a two-step CES demand with elasticities of substitution ρ between retail varieties within sectors, and $\rho > \eta > 1$ between sectors. In this setting, the demand faced by retailer i producing its own variety is:

$$q_{is} = p_{is}^{-\rho} P_s^{\rho-\eta} Q_s \tag{6}$$

where q_{is} is quantity demand of retail variety i, p_{is} is the retailer's price, and Q_s is the sectoral total demand. The sectoral price index is given by $P_s \equiv (\sum_i p_{is}^{1-\rho})^{1/(1-\rho)}$. By definition, market share of firm i is

$$s_{is} \equiv \frac{p_{is} q_{is}}{\sum_{i'} p_{i's} q_{i's}} = (\frac{p_{is}}{P_s})^{1-\rho} \tag{7}$$

Under this setting of the model, demand elasticity of retailer i is then derived[19]

$$\epsilon_{is} \equiv -\frac{d\ln q_{is}}{d \ln p_{is}} = \rho(1 - s_{is}) + \eta s_{is} \tag{8}$$

where s_{is} is retailer i's market share within sector s.

Equation (8) has two important implications as follows: Retail firms with high market shares make their demand less sensitive to their own prices (i.e., low demand elasticities). In addition, high-market-share firms charge higher markups than firms with low market shares. In other words, markups are increasing with market shares, which implies that, unlike Melitz (2003), markup is not constant in this model.

[18] To describe the model, I follow a treatment presented by Amiti et al. (2014), a paper that explores the role of variable markups and imported inputs in low exchange rate pass-through based on Atkeson and Burstein (2008).

[19] Note that $\partial \ln P_s / \partial \ln p_{is} = s_{is}$.

Using (4) - (8), I can determine how retail markups are associated with tariff pass-through. First, the reduction in tariffs give rises to fall in firm's price by Equation (4). When its price decreases, the firm's market share will be higher because the market share is increasing with lower prices, given fixed sectoral price index, as shown by Equation (7). According to Equation (8), the rise in market shares leads to fall in firm's elasticity of demand. As a final step, since tariff reductions make the retail firm increase its markups following decrease in demand elasticity, tariff pass-through go downward due to retail markup channel as I explained in (5). Therefore, it is clear that retail markups have a negative impact on tariff pass-through.

5.2 Retail Market Concentrations and Tariff Pass-Through

Being motivated by the simple theoretical framework described by the previous subsection, I start by estimating the relationship between retail market concentrations and tariff pass-through. In highly concentrated retail markets, retail firms, on average, are able to charge higher markups. It is thus predictable that the impact of tariffs on consumer prices is deteriorated in regions where retail market structure is highly concentrated. To test it, I make use of the Herfindahl index for retail markets which varies across provinces and years. The Herfindahl index measures degrees of retail market concentration, where the index in province r and year t (HI_{rt}) is defined as

$$HI_{rt} = \sum_{f} s_{frt}^2$$

where f denotes a retailer and s_{frt}^2 implies the squared market share of retail firm f in province r and year t. My identification strategy is that I augment the tariff pass-through specification (Equation (1)) by adding the Herfindahl indext (HI_{rt}) and its interaction with tariffs ($HI_{rt} \times \tau_{it}$) on the right hand side. Although HI_{rt} and τ_{it} are respectively sector-specific and province-specific time varying variables, the impact of the interacted variable is able to be estimated by taking advantage of three-way variations in CPI data. The regression specification to explore whether retail market concentrations negatively affect tariff pass-through is the following:

$$\ln p_{irt} = \alpha HI_{rt} + \beta \tau_{it} + \theta (HI_{rt} \times \tau_{it}) + Z_{it}{}' \gamma + X_{rt}{}' \rho + \delta_i + \eta_r + \lambda_t + \varepsilon_{irt} \qquad (9)$$

I include sector(δ_i), province(η_r), year(λ_t) fixed effects and sector-year specific control variables(Z_{it}) as in Equation (1). On top of those controls, I further add province-year specific observable variables(X_{rt}) to alleviate the concern about time varying province-specific confounding components[20]. I expect the estimate for θ to be negative and statistically significant, because the impact of tariffs on consumer prices is formally expressed as

$$\frac{\partial \ln p_{icrt}}{\partial \tau_{it}} = \hat{\beta} + \hat{\theta} HI_{rt} \tag{10}$$

The above equation implies that tariff pass-through varies conditional on retail market concentrations. Suppose that province r has a higher market concentration than province $r^{'}$ (i.e., $HI_{rt} > HI_{r^{'}t} > 0$) and that $\hat{\theta}$ is negative. Then, the tariff pass-through in r will be smaller than in $r^{'}$ as I hypothesize, because

$$\hat{\beta} + \hat{\theta} HI_{rt} < \hat{\beta} + \hat{\theta} HI_{r^{'}t}$$

The key econometric issue when estimating Equation (9) is that the tariff variable is potentially endogenous. In section 4.2, I revealed that the endogeneity is not serious enough to reject the estimation result of Equation (1) itself. In estimating Equation (9), however, it is uncertain whether (and how much degree) the endogeneity contaminates the estimate for θ. Recall that, at least in Equation (1), I mentioned that the feature of Korea's trade reforms constrains us from discovering a proper way to address the endogeneity in tariffs. Meanwhile, Specification (9) can offer an econometric strategy to deal with the issue. To address the endogeneity problem, I focus on the change within sector-year clusters by including sector-year fixed effects (ϕ_{it}) in Equation (9). I expect ϕ_{it} to control for any kinds of sector-year specific confounders certainly including the political economics factor. My preferred regression model for retail markup channel is thus as follows:

$$\ln p_{irt} = \alpha HI_{rt} + \theta (HI_{rt} \times \tau_{it}) + {X_{rt}}^{'} \rho + \phi_{it} + \eta_r + \varepsilon_{irt} \tag{11}$$

I want to note that, compared to Equation (9), τ_{it} and Z_{it} are disappeared in (11) as ϕ_{it} absorbs τ_{it} and Z_{it} as well.

Table 9 illustrates evidence that retail market concentrations matter with tariff pass-through determinations. Columns (1)-(2) report regression results of Equation (9). Column (1) shows results when excluding province-year specific controls(X_{rt}), whereas column (2) presents the one with X_{rt} used. The parameter of interest is θ, which identifies the extent to which market concentration level in the retail sector exacerbates tariff pass-through. In both (1) and (2), the estimates for θ are negative as expected, but is not statistically significant. Column (3) demonstrates estimation results of Equation (11) in which the endogeneity issue

[20]The controls that I use are log population density and log gross regional domestic product (GRDP), both of which are varying across provinces and years.

embodied in tariffs is completely controlled by sector-year fixed effects(ϕ_{it}). The estimate for θ is now statistically significant at the 5 percent level, suggesting that a 10 percentage point higher Herfindahl index causes 0.18 percent lower tariff pass-through at the province level. The result contrasts sharply with columns (1)-(2) where the endogeneity is unsolved, in respects of both significances and magnitudes. Column (1) and (2) have lower $\hat{\theta}$ than column (3) (i.e., $\hat{\theta}$ in columns (1)-(2) has a downward bias), which indicates that political economics factor significantly matters at least to CPI's sensitivity to trade reforms. As a result, Table 9 strongly confirms that the hypothesis "the higher retail market concentration is, the lower tariff pass-through to CPI is." is true.

5.3 The Role of Retail Markups

I focused on the retail market structure that contributes to the incomplete reaction of consumer prices to tariff cuts. Obviously, the rationale behind the argument is that markups charged by retail firms function as market frictions negatively affecting tariff pass-through. The Herfindahl index is, however, not a markup measure but an indicator of how much degree a market structure is concentrated.[21] To reveal more precisely whether the retail channel acts as a friction when tariff cuts pass through to consumer prices, the focus should be on the impact of *retail markups* on CPI sensitivity to tariffs. The corresponding regression specification is, thus, formulated by

$$\ln p_{irt} = \alpha\mu_{rt} + \theta(\mu_{rt} \times \tau_{it}) + X_{rt}{}'\rho + \phi_{it} + \eta_r + \varepsilon_{irt} \tag{12}$$

where μ_{rt} is the markup in province r and year t. As noted in the previous subsection, sector-year fixed effect ϕ_{it} seeks to control for unobserved confounders, which certainly deals with the endogeneity by political economics factor.

There are two challenging issues in examining the role of retail markups. First, it is not straightforward to measure market-specific markups for the retail sector. In fact, a body of recent papers empirically estimates firm-level markups through the structural model estimations. Despite its elegance in terms of being based on solid theoretical backgrounds, the approach is extremely data demanding that it is hard to employ it to this paper. The methodology requires one to make use of firm-level balance sheet data, because it relies on the estimation of output elasticities with respect to variable inputs. Unfortunately, there is no retail firm level data that covers the entire sample period in this paper built on Korea's

[21] In fact, the Herfindahl index is sometimes used as a markup proxy for a particular industry or a geographical market from the literature. In my view, though, it is not because the index is truly a good proxy for markups, but because of lack of proper data.

context. Second, the endogeneity problem arises once I consider *retail markups* as a channel of incomplete pass-through. Previously, I only care about the endogeneity embodied in tariffs in Equation (9) and (11), and the Herfindahl index was treated as an exogenous variable because retail market structure is assumed to be given exogenously as is done in the literature. Yet, it is hard to believe that retail markups are exogenously given. As a result, it is critically important to control for the endogeneity in the markup variable.

Constructing retail markup measures My markup variable is proxied by price-cost margins constructed by accounting data. While price-cost margin has often been criticised (Bresnahan, 1989), I consider price-cost margins as a good measure to proxy retail markups especially under data limitation in Korea. In particular, there is evidence that the measure is eventually almost similar to econometrically estimated markups at least at the aggregated sector or region level. For instance, Siotis (2003) shows a correlation of price-cost margins with sectorally estimated markups is more than 0.8.

Province-level retail markups are constructed based on KOSTAT(Korea's office of national statistics)'s data for the retail sector. The KOSTAT annually collects information on a sample of Korean retail establishments since 1988. The sample survey encompasses a number of variables, especially those included in income statements. To name a few, it includes annual sales, costs of goods sold, operating costs, and so on. There are two drawbacks in KOSTAT's annual survey. First, it does not contain any information that helps to identify establishments, such as establishment(store) names, firm affiliations and addresses. It is thus impossible to use it as a store-level panel data. However, it provides province codes across establishments belonging to the sample. The province codes enable us to build province-level measures for every variable, all of which are able to be utilised as a panel data. Second, although KOSTAT's retail data is de facto a repeated cross-section data because of no information on store identifications, even the restricted store-level data is only accessible for the post-1996 period. KOSTAT's retail data during the pre-1996 period is provided only as aggregated values, i.e., the sum of values across variables. Figure 6 shows that (average) tariff rates decline considerably during the period of 1988 to 1996. To analyse the impact of tariff reductions for years before 1996, I have no choice but to utilise the aggregate values over the whole sample period (1988-2019) so as to establish consistent markup measures over years.

In principle, price-cost margins measure the relative importance of markups to prices. To correctly measure, it is required to exploit marginal cost (c_{rt}) information, which is in fact quite difficult to obtain. As in the literature, I use a roundabout method as defined by the following:

$$\mu_{rt} = \frac{p_{rt} - c_{rt}}{p_{rt}} = \frac{(p_{rt} - c_{rt})q_{rt}}{p_{rt}q_{rt}} = \frac{profit_{rt}}{sale_{rt}}$$

where p_{rt}, c_{rt} and q_{rt} are price, marginal cost and sale quantity in province r and year t, respectively. By accounting conventions, profits are calculated as annual sales less costs of goods sold and operating costs. Therefore, price-cost margins can be further expressed as

$$\mu_{rt} = \frac{profit_{rt}}{sale_{rt}} = \frac{(\mathrm{AS}_{rt} - \mathrm{COGS}_{rt} - \mathrm{OC}_{rt})}{\mathrm{AS}_{rt}} \tag{13}$$

where AS_{rt}, COGS_{rt} and OC_{rt} denote total annual sales($\equiv \sum_s \mathrm{AS}_{srt}$), total costs of goods sold($\equiv \sum_s \mathrm{COGS}_{srt}$), and total operating costs ($\equiv \sum_s \mathrm{OS}_{srt}$) in province r and year t, respectively. As noted above, all aggregated data necessary to calculate the price-cost margin specified in Equation (13) are accessible from KOSTAT's sample survey for the retail sector.

Endogeneity and Instrument Variables In Equation (12), retail markups μ_{rt} are endogenously determined as the following reason: Retail firms tend to take consumer prices into account seriously when deciding where to open stores. To be specific, retailers are more likely to open stores in geographical markets where consumer prices are relatively high, because higher consumer prices would contribute to higher profits. Then, it would happen that more retail stores enter intensively into markets in which consumer prices relatively high. In turn, retail market competition would be higher in markets with high level of consumer prices. As a result, average markups charged by retailers would become lower in those markets, which potentially leads to biased estimates for α as well as θ in Equation (12).

To address the endogeneity issue, I propose two new instrumental variable strategies. The first instrument variable is the province-level Herfindahl index of the *wholesale* sector. As discussed so far, this paper pays attention to horizontal market structure of the retail sector. I highlight that horizontal market concentrations influence average retail markups, which would eventually affect tariff pass-through. On the other hand, retail firms also have a vertical relationship to wholesale firms. The vertical relationship is clearly documented by the standard industrial classifications, mentioning that establishments in the wholesale sector are defined as those who are involved in selling products to retailers. In this sense, retailers' pricing behaviours are arguably determined by the vertical relationship with wholesale firms.

In particular, retail markups are likely to be influenced by market concentrations of the wholesale sector. My data confirms well the relationship between retail markups and wholesale Herfindahl index. I make a plot of (average) retail markups against (average) wholesale Herfindahl index in Figure 8. As clearly shown in the figure, two variables are negatively associated with each other. It may indicate that wholesale market concentrations would adversely influence retailer's markup.[22] While the relationship between two variables is intuitively appealing, theoretical works on retail markups and wholesale market concentrations are scarce. A plausible channel by which retail markups are affected by the vertical relationship is, as suggested by Gaudin (2016), the bilateral bargaining between upstream and downstream firms. Specifically, great wholesale market powers increase their bargaining powers to retailers as sellers of products. Then, retailers inevitably take higher price of inputs and thus their markups are negatively affected. Furthermore, wholesale market concentration is unlikely to have a direct impact on CPI determinations, because wholesalers affect consumer markets only *via retail firms*.[23] Guided by this logic, I construct the Herfindahl index for the wholesale sector using annual sample survey on Korean wholesalers and use it to instrument for retail markups. The KOSTAT does not only undertake store-level sample survey on retail sector, but it also does establishment-level data on the wholesale sector.

The second variable by which I instrument for the retail markup variable μ_{rt} is ten-year lagged populations. Retailing literature has long emphasised the importance of market sizes as an influential determinant of retail store's location and market entry decisions. Figure 9 illustrates location pattern of a particular retail format, large discount stores, as of 2010.[24] Although Figure 9 only shows locations of large discount stores, it would be a good example of how retail stores are geographically distributed. Figure 9 presents that retail stores are concentratedly spotted on regions with large population, such as Seoul, Kyunggi-do and Pusan. Accordingly, these provinces are likely to have higher competition level, and thus are expected to have lower retail markups.

While populations are often considered as an exogenously given variable, I make use of *ten-year lagged* populations, because using long lagged populations would make the variable far more uncorrelated to current consumer prices. To confirm the validity of the proposed instrument, I run a simple regression that relates consumer prices to ten-year lagged population augmented by sector-, province- and year fixed effects plus province-year specific observed controls.[25] Table 10 reports the estimation results of the simple experiment. As

[22] Instead of the analytic Herfindahl index, I use *empirical* index when drawing Figure 8. I obtain it by regressing the analytic Herfindahl index on province- and year fixed effects. The empirical index aims to construct average Herfindahl index which is to a large degree free from province- and year-level heterogeneities issues. More importantly, since the first stage IV regression model includes the two fixed effect terms, it would be a better way to plot retail markups against the empirical wholesale Herfindahl index rather than the analytic index. Nevertheless, I want to note that using the analytic Herfindahl index makes also very clear negative relationship between two variables.

[23] My data also ensures that the IV (the wholesale Herfindahl index) is uncorrelated with CPI. Regressing (log) CPI on the wholesale Herfindahl index and observable controls(X_{rt}) and fixed effects(δ_i, η_r, λ_t), I find that the coefficient of the index is not statistically significant.

[24] I draw Figure 5 using Korea Chainstore Association (KOCA)'s data on large discount stores.

[25] The regression specification is formulated by the following: $\ln p_{irt} = \beta population_{r,t-10} + X_{rt}{}' \gamma + \delta_i + \eta_r + \lambda_t + \varepsilon_{irt}$

expected, the parameter of 10-year lagged population is estimated to be statistically insignificant. Hence, I would expect that long lagged populations have no impact on CPI, but indirectly do only through retail markups.

IV results Retail markup μ_{rt} is endogenous, and thus its interaction with tariffs, $\mu_{rt} \times \tau_{it}$, is potentially endogenous as well. To implement the instrument variable approach, at least two instruments are required to be applied as I have two endogenous variables. Of course, I already proposed two instrumental variables - the Herfindahl index of the wholesale sector and ten-year lagged population - for addressing the endogeneity. The two proposed IVs, yet, are only targeting at μ_{rt}. Put differently, the two variables may not be able to be good and valid instruments for the interaction term $\mu_{rt} \times \tau_{it}$. For the reason, on top of two proposed IVs, I utilise their interactions with a tariff variable as additional instrument variables.[26]

Table 11 presents the results of Equation (12). Column (1) reports OLS estimation results in which IV strategy is not applied. In contrast, column (2) shows the IV estimation results.[27] The estimate of interest in this paper is that of an interaction between retail markups and tariffs which indicates whether and how much retail markups affect CPI's sensitivity to tariffs. The OLS point estimate for θ, a coefficient of $\mu_{rt} \times \tau_{it}$, is of negative sign and is significant at the 5 % level. Even though the OLS result seemingly corresponds well to my hypothesis that retail markups cause fall in tariff pass-through, it should be biased because the endogeneity problem is not solved. IV results in column (2) also report that the point estimate for θ is negative and statistically significant at the 10 % level, but the magnitude (-0.173) becomes much lower than the OLS esimate (-0.0367). The upward bias shown by the OLS result looks obvious when recalling the reason why retail markups are endogenous. I already explained that the variable is endogenous because retail markups affect consumer prices and, at the same time, consumer prices do influence retail markups. The endogeneity issue is, thus, distilled as a reverse causality problem. It is no wonder that the IV point estimate for θ is smaller than the OLS results, because sign of the reverse causal effect (i.e., the effect of consumer prices on retail markups) is negative.

As an alternative IV regression model, I augment Equation (12) by adding province-year fixed effects. The aim of this strategy is basically similar with when I employ sector-year fixed effects to eliminate political economics factor embodied in tariffs. Put simply, I seek to control for endogeneity in markups through province-year fixed effects, because it will be able

[26]One may argue that tariff variable τ_{it} is also endogenous variable. Recall that, however, I showed how sector-year fixed effects (ϕ_{it}) are able to address the endogeneity embodied in τ_{it}. Thus, the endogeneity issue related to the tariff can be controlled for as long as there is sector-year fixed effects(ϕ_{it}) in the IV regression model.

[27]Column (2) as well as (4) are regression results of the second stage IV model. I show the results of the first stage IV estimation in the appendix.

to absorb any unobserved endogenous selection of retail markups into CPI. The approach is formulated by the following:

$$\ln p_{irt} = \theta(\mu_{rt} \times \tau_{it}) + \phi_{it} + \pi_{rt} + \varepsilon_{irt} \tag{14}$$

where π_{rt} denotes province-year fixed effects. Note that since π_{rt} also absorbs observed variables varying across provinces and years, μ_{rt} and X_{rt} are not appeared in Equation (14). I want to emphasize that $\mu_{rt} \times \tau_{it}$ may still be endogenous. While province-year fixed effects π_{rt} deal with endogeneity of μ_{rt}, it could not address that of $\mu_{rt} \times \tau_{it}$ because the interaction term varies at the sector-by-province-by-year dimension. For the reason, I use interactions between proposed IVs (i.e., 10-year lagged populations, wholesale Herfindahl index) and the tariff variable as instrumental variables for $\mu_{rt} \times \tau_{it}$ in Equation (14). Column (3) and (4) in Table 11 present estimation results of Equation (14).

In terms of both magnitudes and significances, estimation results of column (3) and (4) are very similar to column (1) and (2). This suggests that as long as I instrument μ_{rt} and $\mu_{rt} \times \tau_{it}$, the estimate for θ is robust whether the sector-year fixed effect is used(column (4)) or not(column (2)). I report Kleibergen-Paap rk Wald F statistics for the IV results in column (2) and (4). The statistics aim to reveal results of weak IV tests in the case where more than two IVs and clustered standard errors are utilised. The statistics all exceed critical values developed by Stock and Yogo (2005), meaning that there is no concern of weak IV problems.

In conclusion, IV results indicate that the a 10% point increase in retail markups (price-cost margins) causes 1.73% - 1.96% lower tariff pass-through. The regression results so far clearly confirm that retail markups exacerbate the effect of tariffs on consumer prices.

6 Accounting for Heterogeneities and Cross-Section Dependence

6.1 Market Heterogeneities in Retail Markups

In this subsection, I analyse the retail markup channel of which regression model is specified by Equation (12) and (14), focusing on KOSTAT's *census* data for the years of 1996, 2001, 2005 and 2010. In parallel with sample survey on retailers, the KOSTAT has undertaken store-level retail census in those four years. I noted earlier that using year-by-year sample data seems to be inevitable because Korea's tariff reductions has taken place over a long period of time. Nevertheless, it is worth estimating Equations (12) and (14) by limiting to the four census years for two reasons.

Firstly, retail census data makes us overcome the limitation of price-cost margin (PCM) measures. Recall that I construct PCM measures using aggregate (i.e., total sum of values) data provided by KOSTAT's annual sample survey. Accordingly, the reliability of PCM measures crucially depends on whether or not KOSTAT's retail sample represents population well. On the other hand, there is no need to concern about the reliability issue when constructing PCM by census data, because the retail census covers entire population of Korea's retail stores. Regarding the number of observations, retail census amounts to 600,000-770,000 per year whereas retail sample data contains information on 20,000-50,000 stores per year. Secondly, retail census data allows us to address the potential heterogeneity issue in provincial markup measures. While retail sample data only contains province codes as a geographical identifiers, retail census data has narrower geographical indicators such as county codes and town codes.[28] This implies that, when constructing provincial PCM measures relying on census data, it is possible to account for county- or town-level market heterogeneities by using corresponding fixed effects.

The province-level PCM measure was defined as

$$\mu_{rt} = \frac{profit_{rt}}{sale_{rt}} = \frac{\sum_s profit_{rst}}{\sum_s sale_{rst}}$$

where r, s and t index provinces, stores and years respectively, and $profit_{rt}$ and $sale_{rt}$ are total sums of store-level values within a province r. Note that the measure has nothing to do with a statistical mean but it is rather a ratio of aggregate profits to aggregate sales at the province level. Although I use it inevitably for data issue as I mentioned in section 5.3, some may argue that the measure is not precise enough to fully reflect province-level markups. Now, exploiting a richness of store-level census data, I construct a new PCM measures at the province-level as follows:

$$\mu_{rt}^{\text{Census}1} = \sum_{s \in R_r} (weight_{rst} \times \mu_{rst}) \tag{15}$$

where t is now *census* years. μ_{rst} denotes a PCM of store s in province r and census year t. $weight_{rst}$ is a ratio of store-level sales to total provincial sales, which is formulated by $\frac{sale_{rst}}{\sum_s sale_{rst}}$. The new PCM measure $\mu_{rt}^{\text{Census}1}$ is therefore a weighted average PCM of all stores within a province. It is clear that the new PCM $\mu_{rt}^{\text{Census}1}$ is a better measure than the previous one in the sense that the average statistic is derived from a population of Korea's retail

[28]Even from census data, however, it is inaccessible to store identification information (e.g., store names, firm affiliations, etc.) as in annual sample data.

stores.

Even though exploiting store-level variations based on entire population of retail stores, the measure $\mu_{rt}^{\text{Census}1}$ may has another limitation that geographical heterogeneities within a province are not accounted for. Irrespective of locations, the former measures take every store into account identically. In fact, provinces may be too spacious to be a geographical market for a retail store. While I use sales as weights for measuring $\mu_{rt}^{\text{Census}1}$, it may be concerned that market heterogeneities would cause $\mu_{rt}^{\text{Census}1}$ to be incorrectly measured.

Using the census data where detailed geographical codes are available, I further establish a new PCM measure in which county-level market heterogeneity is controlled for. I construct it in three steps. As a first step, I compute weighted average of store-level PCM at the *county* level. The weight is store-level sales proportion as noted before. Thus, the formal specification for weighted average PCM of county j at year t (μ_{rjt}) is

$$\mu_{rjt} = \sum_{s \in J_j} \left(\frac{sale_{rjst}}{\sum_{s \in J_j} sale_{rjst}} \times \mu_{rjst} \right)$$

where r indexes provinces, j denotes counties, s implies stores and t indicates census-years. J_j is a set of store s belonging to county j. The second step is to generate *empirical* county-level PCMs by running a simple regression as follows:

$$\mu_{rjt} = \alpha + \sum_{j=1}^{251} \gamma_j county_j + \sum_{t=1}^{4} \delta_t year_t + \varepsilon_{rjt}$$

The equation above is nothing but a regression of μ_{rjt} on county fixed effects and census-year fixed effects. The aim of the regression is to derive a fitted value of PCM at county j in census year t, which is to a considerable degree free from heterogeneity issues. Let $\hat{\mu}_{rjt}$ denote fitted PCMs obtained by the second step. Then, the last step is to construct the provincial weighted average across fitted county-level PCMs($\hat{\mu}_{rjt}$) as below:

$$\mu_{rt}^{\text{Census}2} = \sum_{j \in R_r} \left(\frac{sale_{rjt}}{\sum_j sale_{rjt}} \times \hat{\mu}_{rjt} \right) \tag{16}$$

I re-estimate Equation (12) and (14), replacing retail markup measures by $\mu_{rt}^{\text{Census}1}$ and $\mu_{rt}^{\text{Census}2}$, and using data for four census years. The instrument variables used in the new estimations are same as before (i.e., ten-year lagged populations and wholesale Herfindahl index for the four census years). Table 12 shows the estimation results of the regressions. Columns (1)-(4) report the results when I use weighted average PCM $\mu_{rt}^{\text{Census}1}$ for the markup measure. Among them, column (1) and (2) are estimation results of Equation (12), whereas column

(3) and (4) are estimation results of Equation (14). Regarding the estimated parameter for the interaction between retail markups and tariffs, OLS results in (1) and (3) are not significant unlike Table 11, but IV results shown in (2) and (4) are significant at the 10% level and negative. Strikingly, the *negative* impact of retail markups on tariff pass-through estimated by IV approach substantially decline compared to those in Table 11. This suggests that accounting for store-level variations in markups by exploiting $\mu_{rt}^{\text{Census}1}$ is meaningful. Nonetheless, the IV results clearly indicate that retail markups negatively affect tariff pass-through even after store-level census data is exploited: 10 percentage point increase in retail markups (price-cost margins) causes 0.64-0.69 percent lower tariff pass-through.

Columns (5)-(8) are estimation results when I employ another markup measure $\mu_{rt}^{\text{Census}2}$, which is weighted average PCM with county-level market heterogeneity controlled. While being slightly different in magnitudes, the estimates reported in columns (5)-(8) show very similar results as when I use $\mu_{rt}^{\text{Census}1}$. The implication of the results is that it is important to account for the market heterogeneity, but once average markups are constructed based on store-level variations, it seems to make little difference. Most importantly, I find that there is a negative causal effect of retail markups on CPI's sensitivity to tariffs, even after controlling for the market heterogeneity in markup measures: 10 percentage point increase in retail markups (price-cost margins) causes 0.85-0.89 percent lower tariff pass-through.[29]

6.2 Cross-Section Dependence in Consumer Prices

This subsection assess the spatial dependence between consumer prices. One may argue that each provincial CPI is affected by consumer prices of others. From a perspective of econometrics, the cross-section dependence would cause a serious identification problem, which results in biased estimates of coefficients in the regression model. To control for it, I exploit an approach in the field of spatial econometrics to take into account interactions across provincial consumer prices.

I run the retail markup channel model specified in Equation (12) and (14), by adopting the spatial autoregressive (SAR) model. More specifically, I include a spatial lag term ($W \ln p_{icrt}$) on the right hand side of my regression model, which is given by

$$\ln p_{irt} = \alpha\mu_{rt} + \theta(\mu_{rt} \times \tau_{it}) + X_{rt}{}'\rho + \phi_{it} + \eta_r + \sigma W \ln p_{icrt} + \varepsilon_{irt} \tag{17}$$

where W denotes a spatial weight matrix of which each element captures the inverse distance

[29]In the Appendix, I report the results when I employ a retail markup measure where, rather than county-specific heterogeneity, town-level market heterogeneity is controlled by the same approach described above. I find that accounting for town-level heterogeneity provides quite robust results.

between a pair of provinces. The assumption behind the approach is that province-pairs with short distances are closely interconnected in terms of consumer price determinations, whereas the effect of remote provinces is not big that much.[30]

I construct a block-diagonal matrix for W. Each block, w, represents a $P \times P$ matrix for each sector-year group, where P is the number of provinces. The $P \times P$ matrix is defined as

$$w = \begin{bmatrix} 0 & {d_{12}}^{-1} & {d_{13}}^{-1} & \dots & {d_{1P}}^{-1} \\ {d_{21}}^{-1} & 0 & {d_{23}}^{-1} & \dots & {d_{2P}}^{-1} \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ {d_{P1}}^{-1} & {d_{P2}}^{-1} & {d_{P3}}^{-1} & \dots & 0 \end{bmatrix}$$

where every off-diagonal element is an inverse of the distance from a particular province p to another province p'. I normalize the matrix w by min-max normalizing procedure.[31] The full weight matrix W is an $SPT \times SPT$ matrix where S is total number of sectors and T is total number of years. The full weight matrix which is a block-diagonal matrix given by

$$W = \begin{bmatrix} w & 0 & 0 & 0 \\ 0 & w & 0 & 0 \\ 0 & 0 & \ddots & 0 \\ 0 & 0 & 0 & w \end{bmatrix}$$

Table 13 reports estimation results of Equation (17). Even incorporating spatial autoregressive (SAR) model, the results show simliar pattern with those in Table 11 where SAR model is not included. In particular, IV results in Table 13 suggest that increase in retail markups by 10 percentage points causes 1.10-1.12 percent lower tariff pass-through. The negative effect of retail markups on tariffs slightly declines compared to the regression without SAR model, which indicates that cross-section dependence exists spatially. Nevertheless, the SAR model makes sure that even controlling for the cross-section yields strong evidence that retail markups matter for forming imperfect tariff pass-through.

[30] Another way to account for interactions across province-level CPI is spatial error model(SEM), where the structure of spatial interaction term is put on the errors. I choose SAR model rather than SEM, because I already make use of standard errors clustered at the sector-year level throughout this paper. I expect that the clustered standard errors will greatly alleviate the concern about cross dependences across regions in the *error* term.

[31] The min-max normalized matrix is as the following: the $\{p,\ p'\}$th element of the normalized matrix becomes the value of the original element divided by m, where $m = \min\{$the largest row sum of pre-normalized matrix, the largest column sum of pre-normalized matrix$\}$.

7 Conclusion

I provide evidence on the relationship between tariffs and consumer prices. I show that consumer prices are incompletely responding to tariff reductions. More importantly, I find that the imperfect tariff pass-through is crucially driven by the markups charged by retail intermediaries. While the retail markup channel is investigated by a few exchange rate pass-through literature, it has rarely been studied in the context of trade reforms. According to Bernard et al. (2010), intermediary firms "affect the magnitude and nature of trade frictions and both the pattern of trade and its welfare gains". The findings in this paper strongly suggest that retail intermediaries play a crucial role in determining how much degree consumers benefit from trade liberalisation. Again, this paper relates to trade induced macroeconomic issues such as welfare analysis, inequality and poverty, which can be explored by future research.

References

Amiti, M., Itskhoki, O. and Konings, J. (2014), 'Importers, exporters, and exchange rate disconnect', *American Economic Review* **104**(7), 1942–78.

Atkeson, A. and Burstein, A. (2008), 'Pricing-to-market, trade costs, and international relative prices', *American Economic Review* **98**(5), 1998–2031.

Atkin, D. and Donaldson, D. (2015), Who's getting globalized? the size and implications of intra-national trade costs, Technical report, National Bureau of Economic Research.

Bernard, A. B., Jensen, J. B., Redding, S. J. and Schott, P. K. (2010), 'Wholesalers and retailers in us trade', *The American Economic Review* pp. 408–413.

Bresnahan, T. F. (1989), 'Empirical studies of industries with market power', *Handbook of industrial organization* **2**, 1011–1057.

Broda, C. and Romalis, J. (2008), Inequality and prices: Does china benefit the poor in america, Technical report, University of Chicago (mimeo).

Burstein, A. and Gopinath, G. (2014), International prices and exchange rates, *in* 'Handbook of international economics', Vol. 4, Elsevier, pp. 391–451.

Burstein, A. T., Neves, J. C. and Rebelo, S. (2003), 'Distribution costs and real exchange rate dynamics during exchange-rate-based stabilizations', *Journal of monetary Economics* **50**(6), 1189–1214.

Cheung, J. and Hongsik, L. (2010), 'Changes in korean tariff structure under the new free trade agreements', *Korean Institute of Public Finance* .

Cho, J., Chun, H. and Lee, Y. (2015), 'How does the entry of large discount stores increase retail employment? evidence from korea', *Journal of Comparative Economics* **43**(3), 559–574.

Corsetti, G. and Dedola, L. (2005), 'A macroeconomic model of international price discrimination', *Journal of International Economics* **67**(1), 129–155.

De Loecker, J., Goldberg, P. K., Khandelwal, A. K. and Pavcnik, N. (2016), 'Prices, markups, and trade reform', *Econometrica* **84**(2), 445–510.

De Loecker, J. and Warzynski, F. (2012), 'Markups and firm-level export status', *American economic review* **102**(6), 2437–71.

Dornbusch, R. (1987), 'Exchange rates and prices', *American Economic Review* **77**(1), 93–106.

Faber, B. (2014), Trade liberalization, the price of quality, and inequality: Evidence from mexican store prices, Technical report, Mimeo, University of California at Berkeley.

Feenstra, R. C. (1989), 'Symmetric pass-through of tariffs and exchange rates under imperfect competition: An empirical test', *Journal of international Economics* **27**(1), 25–45.

Gaudin, G. (2016), 'Pass-through, vertical contracts, and bargains', *Economics Letters* **139**, 1–4.

Goldberg, L. S. and Campa, J. M. (2010), 'The sensitivity of the cpi to exchange rates: Distribution margins, imported inputs, and trade exposure', *The Review of Economics and Statistics* **92**(2), 392–407.

Han, J., Liu, R., Marchand, B. U. and Zhang, J. (2016), 'Market structure, imperfect tariff pass-through, and household welfare in urban china', *Journal of International Economics* **100**, 220–232.

Hong, G. H. and Li, N. (2017), 'Market structure and cost pass-through in retail', *Review of Economics and Statistics* **99**(1), 151–166.

Kim, J.-I. and Kim, J.-D. (2003), Liberalization of trade in services and productivity growth in korea, *in* 'Trade in Services in the Asia Pacific Region, NBER East Asia Seminar on Economics (EASE), Volume 11', University of Chicago Press, pp. 179–208.

Marchand, B. U. (2012), 'Tariff pass-through and the distributional effects of trade liberalization', *Journal of Development Economics* **99**(2), 265–281.

Melitz, M. J. (2003), 'The impact of trade on intra-industry reallocations and aggregate industry productivity', *Econometrica* **71**(6), 1695–1725.

Nakamura, E. (2008), 'Pass-through in retail and wholesale', *American Economic Review* **98**(2), 430–37.

Nakamura, E. and Zerom, D. (2010), 'Accounting for incomplete pass-through', *The Review of Economic Studies* **77**(3), 1192–1230.

Nicita, A. (2009), 'The price effect of tariff liberalization: Measuring the impact on household welfare', *Journal of Development Economics* **89**(1), 19–27.

Oster, E. (2017), 'Unobservable selection and coefficient stability: Theory and evidence', *Journal of Business & Economic Statistics* pp. 1–18.

Porto, G. G. (2006), 'Using survey data to assess the distributional effects of trade policy', *Journal of International Economics* **70**(1), 140–160.

Siotis, G. (2003), 'Competitive pressure and economic integration: An illustration for spain, 1983–1996', *International Journal of Industrial Organization* **21**(10), 1435–1459.

Spengler, J. J. (1950), 'Vertical integration and antitrust policy', *Journal of political economy* **58**(4), 347–352.

Stock, J. H. and Yogo, M. (2005), 'Testing for weak instruments in linear iv regression'.

Topalova, P. (2010), 'Factor immobility and regional impacts of trade liberalization: Evidence on poverty from india', *American Economic Journal: Applied Economics* **2**(4), 1–41.

Table 1: Average Agricultural Tariff Rates

Year	1989	1990	1992	1995	1996	1997	2000	2005	2010
Tariff (%)	16.71	22.39	23.09	21.98	69.88	73.35	64.72	50.72	53.22

Notes: This table shows average agricultural tariff rates over years. To calculate them, I use UNCTAD-TRAINS and WTO-IDB data. As noted in section 3.1, there are missing values in tariff rates for the years of 1991, 1993 and 1994, even in the combination of two sources of data. That is the reason why this table is not including values for the three years.

Table 2: Target Average Tariff Rates

Year	1988	1989	1990	1991	1992	1993	1994
Tariff (%)	20.0	15.0	13.0	12.0	11.0	9.0	8.0

Source: Cheung and Hongsik (2010)

Notes: This table presents the average target tariff rates set by the Korean government. Korea carried out tariff reductions in a manner that it followed pre-determined reduction plan, and as shown in this table, the target rates are clearly monotonically decreasing trend.

Table 3: Descriptive Statistics

Variable Name	N	Mean	St. Dev.	Min	Max
Log CPI ($CPI_{2010} = 100$)	9,750	4.500	0.531	2.670	6.881
Weighted Average Tariff Rates (%)	9,984	7.975	4.1171	0.000	23.201
Herfindahl Index	6,630	0.031	0.028	0.001	0.182
Log World Price ($)	9,152	4.418	2.967	0.000	10.724
Import Penetration Ratio (%)	9,568	0.021	0.038	0.000	0.244
Log imports from China to Korea	9,616	11.322	2.294	1.971	15.597
Log GRDP (Million KRW)	10,114	17.415	0.848	15.050	19.534
Log Population Density (per km^2)	10,140	6.619	1.545	4.489	9.759

Notes: This table shows summary statistics for the main variables used in the empirical analyses of this paper.

Table 4: Sector-level Log Consumer Price Indices over Years

	1990		1995		2000		2005		2012	
Sector Name	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Books	3.71	0.03	3.97	0.01	4.28	0.01	4.44	0.00	4.67	0.00
Casual Clothes	3.96	0.08	4.20	0.00	4.27	0.00	4.41	0.00	4.69	0.00
Children's Clothes	4.02	0.16	4.24	0.04	4.37	0.00	4.55	0.00	4.71	0.00
Electric Applicances for Personal Cares	3.59	0.08	4.06	0.09	4.35	0.05	4.45	0.03	4.71	0.02
Equipment for Sound and Pitures	5.95	0.07	5.83	0.07	5.58	0.05	5.31	0.03	4.36	0.02
Fuels and Lubricants for transport Equipment	3.03	0.03	3.44	0.02	4.20	0.02	4.37	0.01	4.76	0.01
Furnitures, Furnishing and Carpets	3.88	0.06	4.10	0.05	4.17	0.03	4.38	0.00	4.68	0.00
Glassware, Tableware and Household utensils	3.97	0.07	4.24	0.06	4.36	0.04	4.46	0.02	4.72	0.02
Household Appliances	4.93	0.07	4.87	0.06	4.73	0.05	4.62	0.03	4.61	0.01
Household Textile	4.24	0.12	4.43	0.07	4.53	0.04	4.43	0.06	4.70	0.01
Information Processing Equipments	6.43	0.12	6.08	0.06	5.65	0.08	5.20	0.02	4.46	0.01
Medical Products, appliances and Equipments	4.20	0.09	4.42	0.05	4.48	0.05	4.57	0.02	4.64	0.01
Men's Clothes	4.17	0.04	4.29	0.02	4.38	0.01	4.42	0.01	4.70	0.01
Non-Durable Houshold Goods	3.94	0.08	4.10	0.06	4.32	0.06	4.44	0.04	4.70	0.03
Other Articles of Clothing and Accessories	4.25	0.06	4.36	0.05	4.49	0.02	4.52	0.02	4.67	0.01
Othe Fuels and Energy	2.90	0.10	3.14	0.07	3.92	0.07	4.37	0.05	4.82	0.04
Othe Major Durables for Recreation and Culture	3.92	0.06	4.20	0.04	4.37	0.04	4.49	0.04	4.65	0.02
Other Recreational Items, Gardens and Pets	4.09	0.06	4.30	0.06	4.45	0.03	4.48	0.02	4.65	0.01
Personal Effects N.E.C.	4.28	0.06	4.23	0.06	4.37	0.04	4.42	0.01	4.64	0.00
Pharmaceutical Products	4.33	0.33	4.50	0.04	4.47	0.02	4.57	0.01	4.61	0.02
Purchase of Vehicles	4.53	0.01	4.55	0.00	4.53	0.00	4.54	0.00	4.61	0.00
Shoes and Other Footwear	3.94	0.05	4.05	0.02	4.28	0.01	4.51	0.01	4.69	0.01
Stationary and Drawing Materials	3.82	0.09	4.15	0.06	4.41	0.04	4.39	0.02	4.72	0.02
Telephone and Telefax Equipments	6.54	0.12	6.50	0.09	6.07	0.07	5.20	0.01	4.52	0.01
Tools and Equipment for House and Garden	4.63	0.16	4.63	0.11	4.53	0.07	4.49	0.05	4.65	0.02
Women's Clothes	4.05	0.02	4.30	0.01	4.36	0.01	4.44	0.01	4.67	0.00

Notes: This table reports sector-level log CPI over Years. A sector-level log CPI is calculated by averaging out all provincial log CPI within each sector. A sector-level standard deviation of log CPI measures the degree to which province-level indices are dispersed within each sector.

Table 5: Province-level Log Consumer Price Indices over Years

	1990		1995		2000		2005		2012	
Province Name	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Cheju	4.30	0.88	4.44	0.74	4.54	0.52	4.55	0.27	4.65	0.10
Chungbuk	4.31	0.86	4.45	0.74	4.54	0.50	4.55	0.26	4.65	0.09
Chungnam	4.26	0.85	4.44	0.72	4.55	0.51	4.56	0.26	4.65	0.09
Deagion	4.25	0.88	4.45	0.72	4.56	0.49	4.56	0.26	4.66	0.09
Deagu	4.33	0.92	4.45	0.74	4.54	0.49	4.57	0.27	4.65	0.09
Gangwon	4.28	0.85	4.43	0.72	4.53	0.48	4.55	0.27	4.65	0.09
Gwangju	4.28	0.85	4.41	0.69	4.52	0.45	4.55	0.26	4.65	0.09
Incheon	4.27	0.84	4.43	0.71	4.53	0.48	4.56	0.26	4.65	0.09
Junbuk	4.26	0.88	4.42	0.71	4.53	0.47	4.56	0.26	4.66	0.10
Junnam	4.28	0.86	4.42	0.72	4.54	0.49	4.55	0.26	4.66	0.10
Kyungbuk	4.29	0.85	4.43	0.71	4.53	0.48	4.55	0.26	4.66	0.09
Kyunggi	4.26	0.86	4.42	0.72	4.53	0.49	4.55	0.26	4.65	0.09
Kyungnam	4.28	0.85	4.42	0.72	4.53	0.48	4.54	0.26	4.65	0.10
Pusan	4.26	0.84	4.42	0.71	4.53	0.46	4.55	0.25	4.66	0.09
Seoul	4.30	0.88	4.43	0.74	4.54	0.49	4.54	0.27	4.65	0.09
Woolsan			4.42	0.72	4.54	0.48	4.55	0.26	4.66	0.09

Notes: This table reports province-level log CPI over Years. Each province-level index is the avarage of sectoral log consumer price indices within the province. Standard deviations present degree of variations in log CPI at the province level.

Table 6: Tariff Rates and Consumer Price Index

	Dependent variable: Log CPI				
	(1)	(2)	(3)	(4)	(5)
Tariff Rates	-0.000962	-0.000621	-0.000181	0.0656***	0.0599***
	(0.00501)	(0.00782)	(0.00783)	(0.00996)	(0.00965)
Import Penetration Ratio					3.252***
					(0.596)
Log World Price					0.0775***
					(0.0144)
Log Imports from China					-0.153***
					(0.0195)
Sector FE	No	Yes	Yes	Yes	Yes
Province FE	No	No	Yes	Yes	Yes
Year FE	No	No	No	Yes	Yes
N	9,594	9,594	9,594	9,594	8,716
R^2	0.000	0.612	0.613	0.689	0.745
Clusters	624	624	624	624	551

Notes: This table demonstrates the association between tariffs and consumer prices. *Import Penetration Ratio* measures sector-level degrees of import penetration defined as a ratio between imports and imports plus domestic productions at the sector level. (Log) *World Price* varies across sectors and years and it is proxied by U.S.'s average export prices of products within each sector-year cluster. * p-value < 0.10, ** p-value < 0.05, *** p-value < 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

Table 7: Assessment of Bias Due to Unobserved Confounding Effects

	Regression Results Without Controls (1)	Regression Results With Observables (2)
Tariff Rates	-0.000962 (β^1)	0.0599 (β^2)
R^2	0.0001 (${R^2}_1$)	0.7453 (${R^2}_2$)
Bias due to Unobserved Confouding Effects	-0.0208	

Notes: This table shows a result of sensitivity analysis based on Oster (2017)'s approach. Column (1) reports the regression results when I regress log CPI on tariffs without any control variables. On the other hand, column (2) displays regression results when using observable controls. The observables exploited for column (2) are sector fixed effect, province fixed effect, year fixed effect and three sector-year specific control variables described in section 4.1. I present the bias driven by unobserved confounding effects at the bottom of the table. It is calculated by the formula below,

$$\text{Confounding Effect} = \delta \frac{(\beta^1 - \beta^2)({R^2}_{max} - {R^2}_2)}{({R^2}_2 - {R^2}_1)}$$

To calculate it, I use regression results reported in columns (1)-(2). As noted in section 4.1, δ and ${R^2}_{max}$ are all assumed to be one, because the value is the hardest case to happen.

Table 8: Log(1+Tariff) and CPI

	Dependent variable: Log CPI				
	(1)	(2)	(3)	(4)	(5)
Tariff Rates	0.0175	-0.0398	-0.0377	0.217***	0.212***
	(0.0237)	(0.0624)	(0.0625)	(0.0728)	(0.0653)
Import Penetration Ratio					3.101***
					(0.625)
Log World Price					0.0646***
					(0.0134)
Log Imports from China					-0.170***
					(0.0204)
Sector FE	No	Yes	Yes	Yes	Yes
Province FE	No	No	Yes	Yes	Yes
Year FE	No	No	No	Yes	Yes
N	9,594	9,594	9,594	9,594	8,716
R^2	0.001	0.613	0.614	0.663	0.728
Clusters	624	624	624	624	551

Notes: This table reports the estimation results of Equation (2) where $\ln(1+\tau_{it})$ is used as an independent variable instead of τ_{it}. Results are almost identical to those in the original model (Table 6). *Import Penetration Ratio* measures sector-level degrees of import penetration defined as ratios between imports and imports plus domestic productions. (Log) *World Price* varies across sectors and years and it is proxied by U.S.'s average export prices across products within each sector-year cluster. * p-value $<$ 0.10, ** p-value $<$ 0.05, *** p-value $<$ 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

Table 9: Retail Market Concentration and Tariff Pass-Through

	Dependent variable: Log CPI		
	(1)	(2)	(3)
Herfindahl Index	0.454	0.448	0.153*
	(0.601)	(0.601)	(0.0789)
Tariff Rates	0.0521***	0.0521***	
	(0.0167)	(0.0167)	
Herfindahl Index × Tariff Rates	-0.0588	-0.0588	-0.0180**
	(0.0818)	(0.0818)	(0.00715)
Sector-Year Observables	Yes	Yes	No
Province-Year Observables	No	Yes	Yes
Sector FE	Yes	Yes	No
Province FE	Yes	Yes	Yes
Year FE	Yes	Yes	No
Sector-Year FE	No	No	Yes
N	6,390	6,390	6,390
R^2	0.662	0.662	0.993
Clusters	401	401	401

Notes: This table presents that increase in retail market concentration measured by the Herfindahl index lowers tariff pass-through to consumer prices. In particular, the regression result shown by column (3), where endogeneity embodied in tariffs is controlled by *sector-year fixed effects* provide evidence of the negative causal effect of retail market concentrations on the pass-through. Sector-year specific observable controls used for the regressions are import penetration ratio, (log) world prices and (log) imports from China. Province-year observables are (log) population density and (log) gross regional domestic product. * p-value < 0.10, ** p-value < 0.05, *** p-value < 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

Table 10: Ten-Year Lagged Populations and Consumer Prices

	Dependent variable: Log CPI	
	(1)	(2)
Population_{t-10}	0.000287	-0.000302
	(0.000954)	(0.00119)
Province-Year Observables	No	Yes
Sector FE	Yes	Yes
Province FE	Yes	Yes
Year FE	Yes	Yes
N	9,750	9,646
R^2	0.653	0.650
Clusters	650	650

Notes: This table demonstrates that ten-year lagged population(Population_{t-10}) is not associated with current consumer prices. The results are robust whether province-year specific observables (log population density and log gross regional domestic product) are used (column (2)) or not (column (1)). * p-value $<$ 0.10, ** p-value $<$ 0.05, *** p-value $<$ 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

Table 11: Retail Markup Channel

	Dependent variable: Log CPI			
	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
Retail Markup	0.161	1.985		
	(0.115)	(1.405)		
Retail Markup × Tariff	-0.0367**	-0.173*	-0.0281**	-0.196**
	(0.0142)	(0.0954)	(0.0122)	(0.0966)
Province-Year Observables	Yes	Yes	No	No
Sector-Year FE	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	No	No
Province-Year FE	No	No	Yes	Yes
N	5,550	4,800	5,500	4,800
R^2	0.992	0.992	0.992	0.992
Clusters	350	300	350	300
Kleibergen-Paap rk Wald F		41.685		313.304

Notes: This table illustrates that increase in provincial retail markups lead to fall in tariff pass-through to consumer prices. Column (1) and (3) are OLS results without IVs, whereas Column (2) and (4) shows estimation results of (the second stage) IV regressions. Kleibergen-Paap rk Wald F statistics present the results of week IV test in the case where more than two IVs and clustered standard errors are exploited. In column (2) and (4), the statistics all exceed critical values proposed by Stock and Yogo (2005), which implies that there is no week IV problem in the IV results. Province-year observables are (log) population density and (log) gross regional domestic product. * p-value < 0.10, ** p-value < 0.05, *** p-value < 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

Table 12: Retail Markup Channel, *Analysed by Census Data*

	Weighted Average PCM				*Weighted Average PCM +Controlling for County Heterogeneity*			
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Retail Markup	0.136	0.553			0.0855	0.728*		
	(0.131)	(0.360)			(0.0521)	(0.390)		
Retail Markup × Tariff	-0.0171	-0.0643*	-0.0175	-0.0693*	-0.0062	-0.0853*	-0.0052	-0.0892*
	(0.0134)	(0.0363)	(0.0135)	(0.0365)	(0.0062)	(0.0474)	(0.0060)	(0.0478)
Province-Year Observables	Yes	Yes	No	No	Yes	Yes	No	No
Sector-Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	No	No	Yes	Yes	No	No
Province-Year FE	No	No	Yes	Yes	No	No	Yes	Yes
N	1,590	1,590	1,590	1,590	1,590	1,590	1,590	1,590
R^2	0.994	0.994	0.994	0.994	0.994	0.994	0.994	0.994
Clusters	101	101	101	101	101	101	101	101
Kleibergen-Paap rk Wald F		50.698		490.822		284.164		3012.51

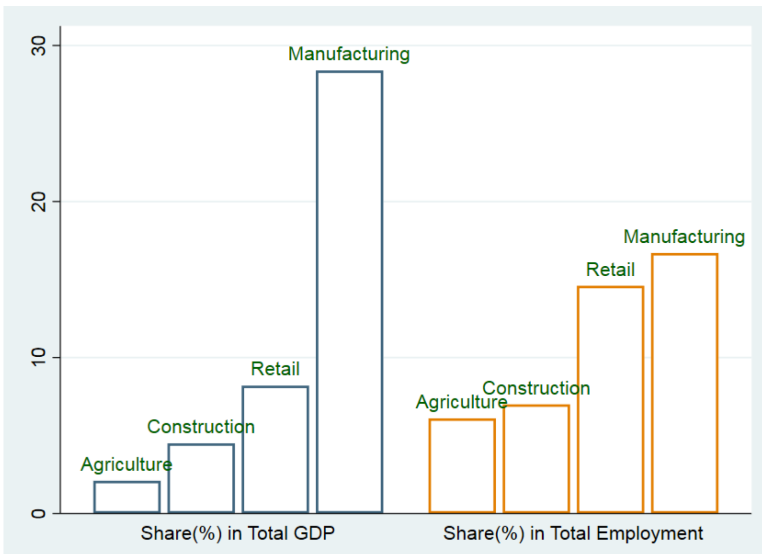
Notes: This table illustrates the IV results which show that increase in provincial retail markups lead to fall in tariff pass-through to consumer prices using the retail census data. Columns (1)-(4) show estimation results when I use weighted average of store-level PCMs ($\mu_{rt}^{\texttt{Census1}}$). On the other hand, columns (5)-(8) report estimation results when, controlling for county-level market heterogeneity, I use weighted average of store-level PCMs ($\mu_{rt}^{\texttt{Census2}}$). Column (1), (3), (5) and (7) are OLS results without IVs, whereas Column (2), (4), (6) and (8) show estimation results of (the second stage) IV regressions. Kleibergen-Paap rk Wald F statistics presents the results of week IV test in the case where more than two IVs and clustered standard errors are exploited. In column (2), (4), (6) and (8), the statistics all exceed critical values proposed by Stock and Yogo (2005), which implies that there is no week IV problem in the IV results. Province-year observables are (log) population density and (log) gross regional domestic product. * p-value < 0.10, ** p-value < 0.05, *** p-value < 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

Table 13: Retail Markup Channel, *Spatial Autoregressive (SAR) Model*

	Dependent variable: Log CPI			
	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
Retail Markup	0.124	0.773^{***}		
	(0.0938)	(0.180)		
Retail Markup × Tariff	-0.0238^{**}	-0.112^{***}	-0.0202^{*}	-0.110^{***}
	(0.0114)	(0.0218)	(0.0114)	(0.0216)
Spatial Lag Term	Yes	Yes	Yes	Yes
Province-Year Observables	Yes	Yes	No	No
Sector-Year FE	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	No	No
Province-Year FE	No	No	Yes	Yes
N	4,992	4,992	4,992	4,992

Notes: This table evaluates whether accounting for spatially cross-section dependence would lead to different estimation results of retail markup channel regression model (Equation (12) and (14)). I find that all the results are similar to Table 11, indicating that even controlling for the cross section dependence yields clear evidence that retail markups play a role in decrease in tariff pass-through. The technique by which the issue of cross-section dependence is dealt with is Spatial Autoregressive (SAR) model. Column (1) and (3) are OLS results without IVs, whereas Column (2) and (4) show estimation results of (the second stage) IV regressions. Province-year observables are (log) population density and (log) gross regional domestic product. * p-value < 0.10, ** p-value < 0.05, *** p-value < 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

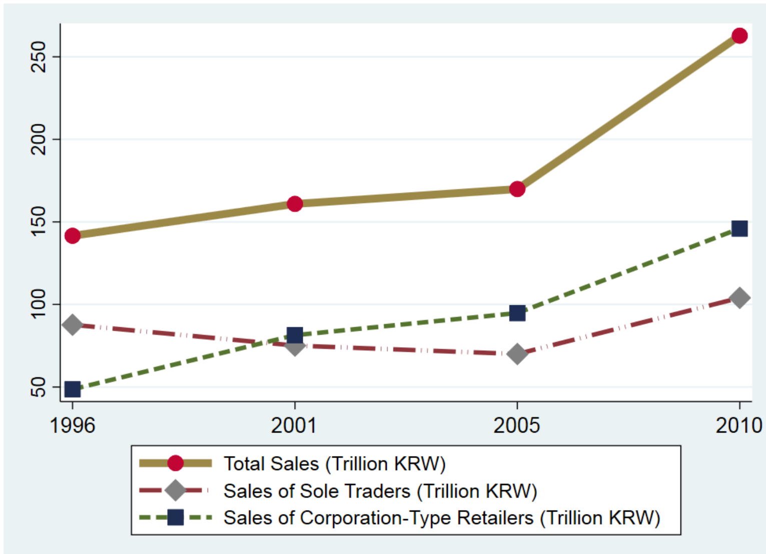
Figure 1: The importance of the retail sector in Korea's economy (2013)



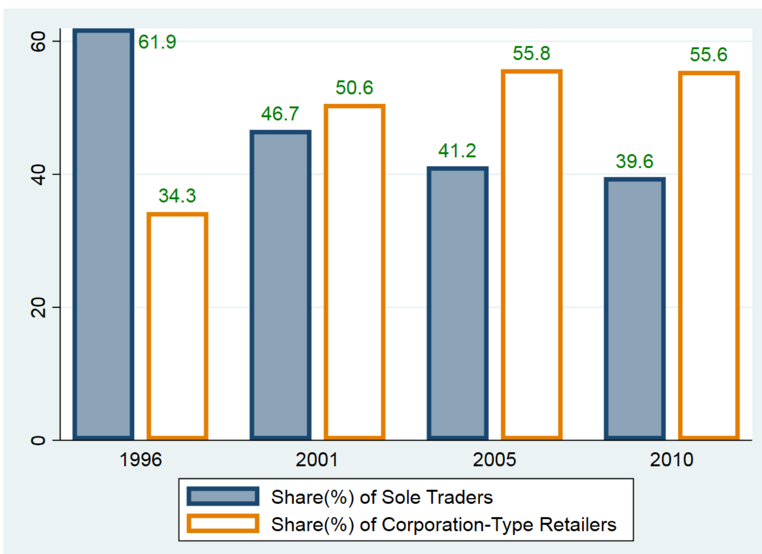
Sources: Bank of Korea, National Account; KOSTAT

Notes: This figure shows the importance of the retail sector relative to other sectors. The left four bars illustrate sectoral GDP shares(%), whereas the four bars on the right hand demonstrate sectoral employment shares(%). All the values are based on statistics in 2013.

Figure 2: The Evolutions of Total Retail Sales



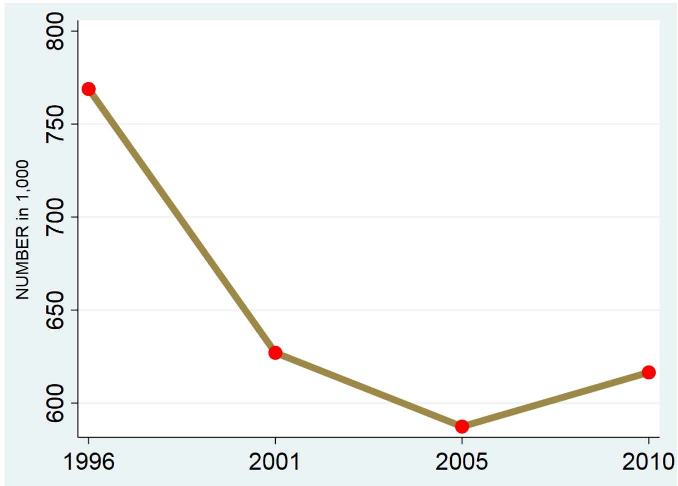
(a) Time Trends in Total Sales



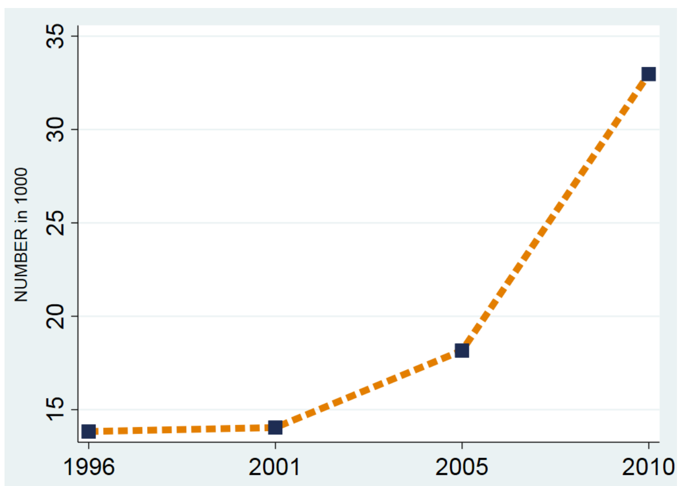
(b) Time Trends in Sales Shares by Business Type

Notes: Panel (a) demonstrates the evolutions of total sales. I further show trends in total sales of sole traders (i.e., self- and family-owned retailers) as well as those of corporation-type retailers. Panel (b) presents sales shares of both sole traders and corporation-type retailers over years. Sales amounts in Panel (a) are expressed in trillion Korean won (KRW) and are deflated by consumer price index.

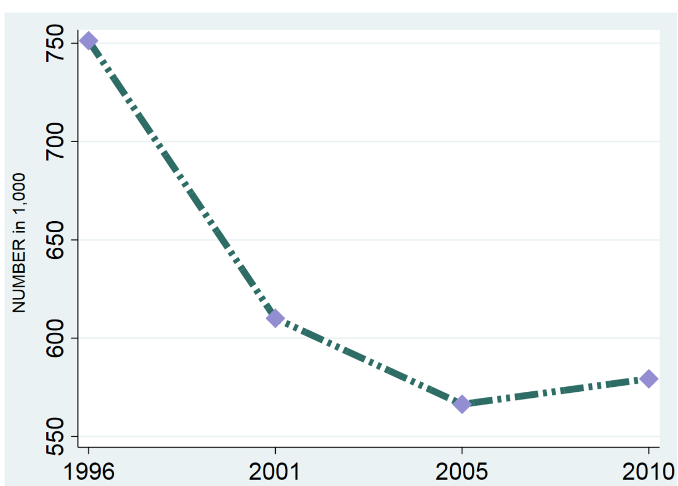
Figure 3: The Evolution of Total Number of Retailers



(a) Total Number of Retailers



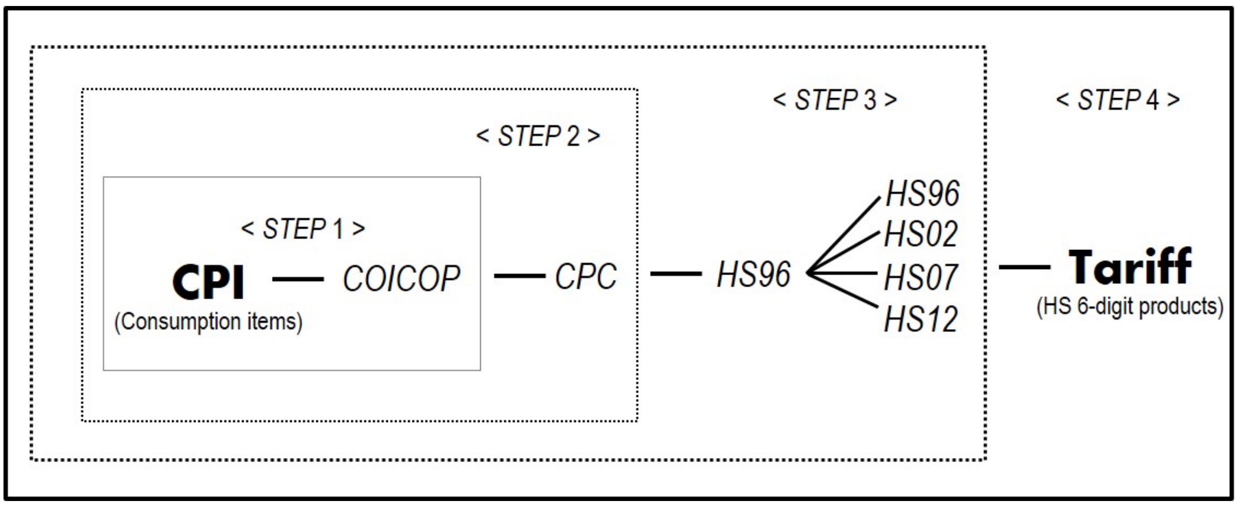
(b) Total Number of Retail Corporations



(c) Total Number of Sole Traders

Notes: Panel (a) presents total numbers of retailers, Panel (b) shows total numbers of corporation-type retailers, and Panel (c) illustrates total number of sole traders. We plot them during the period of 1996 to 2010.

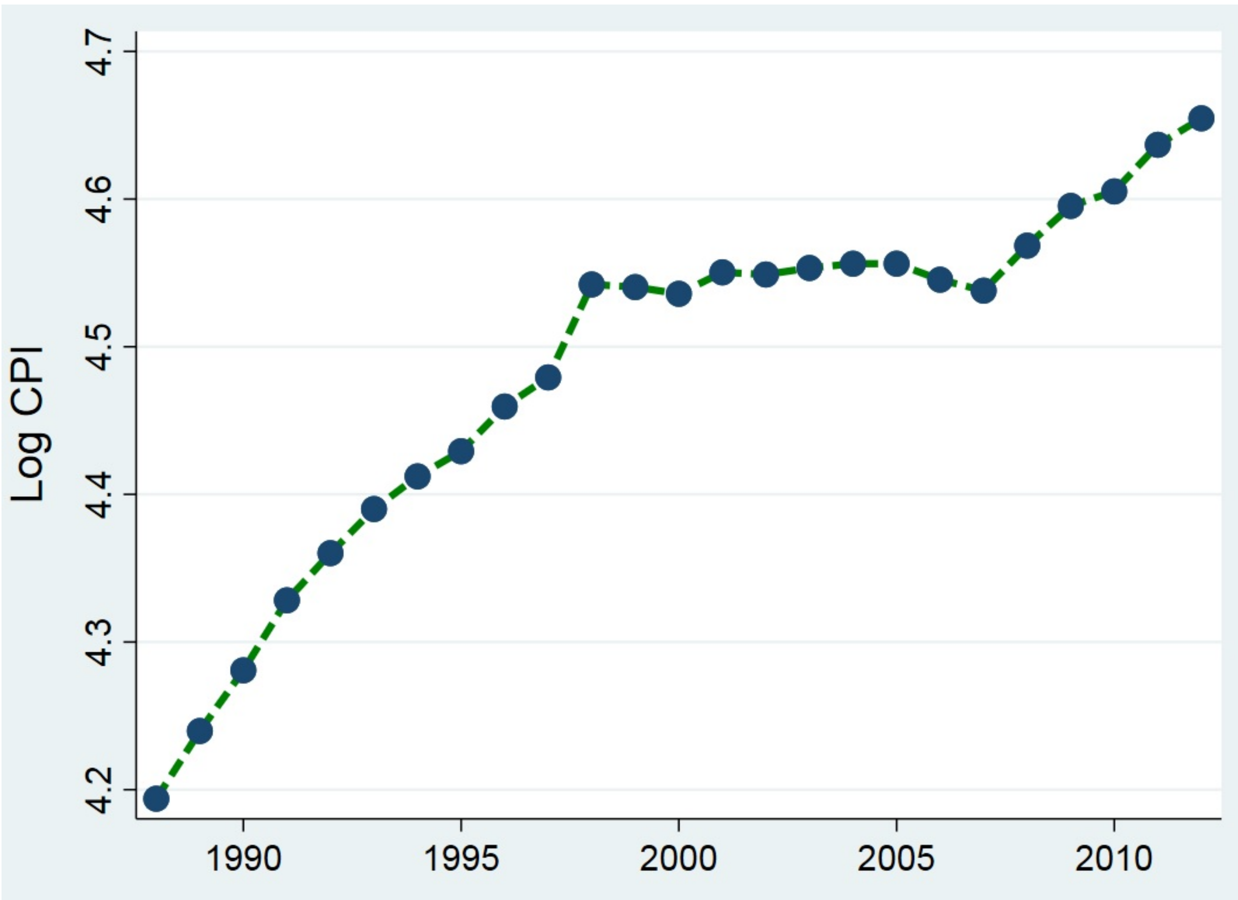
Figure 4: Matching procedure between CPI and Tariffs



Note A: This figure illustrates how I link consumption-item level CPI data to product-level tariff data, which consists of four steps. First, I relate CPI to COICOP (step 1). Then, I merge the matched data with CPC via UNSD's concordance (step 2). I further merge the matched CPC from step 2 to HS codes using both the concordance between CPC and HS nomenclature 1996 and the concordance among different versions of HS codes (step 3). Finally, I construct weighted average tariff rates of all matched products within each CPI sector.

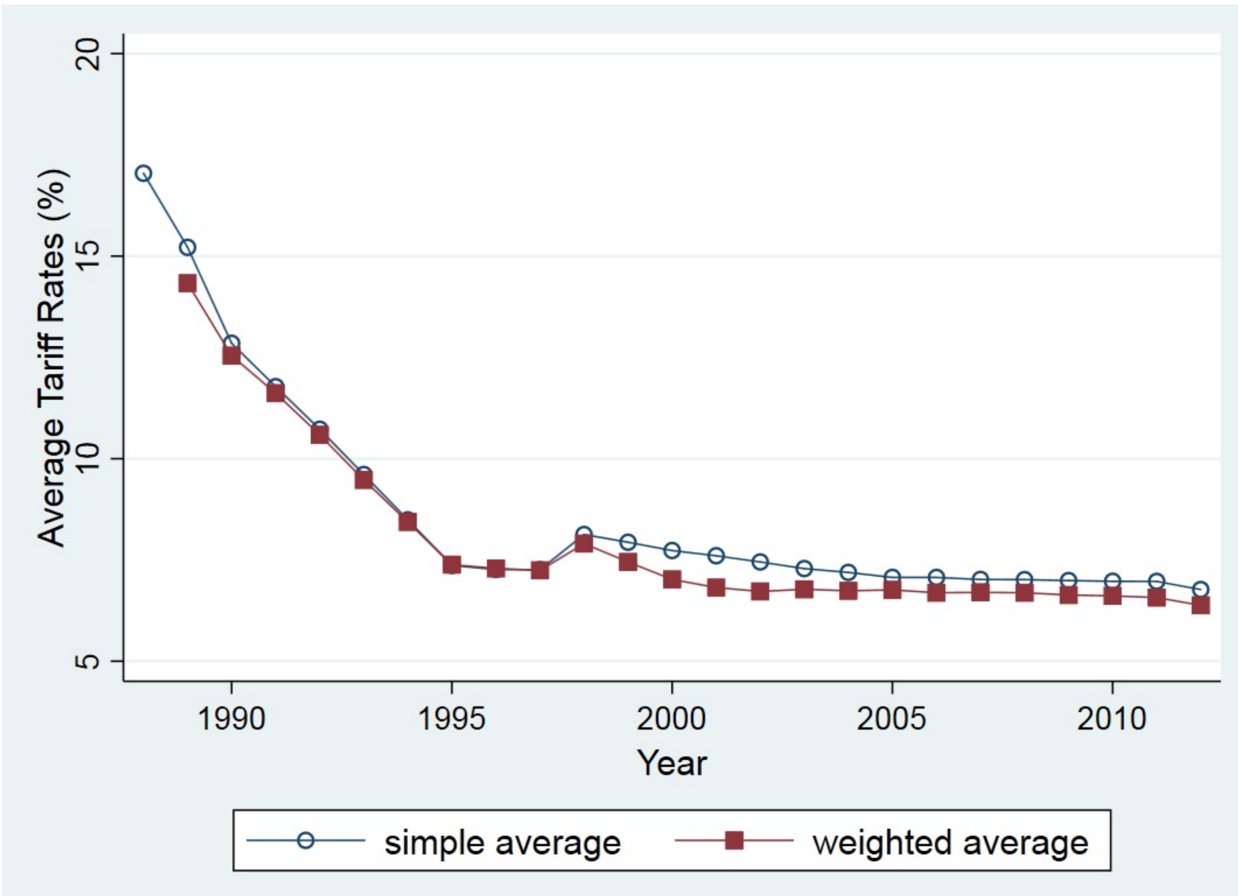
Note B: COICOP is a reference classification for consumption items. CPC is a product classification based on physical characteristics. All the concordances can be found in UNSD(United Nation Statistics Division) classification registry (http://unstats.un.org/unsd/cr/registry).

Figure 5: Time Trend in Log Consumer Price Index



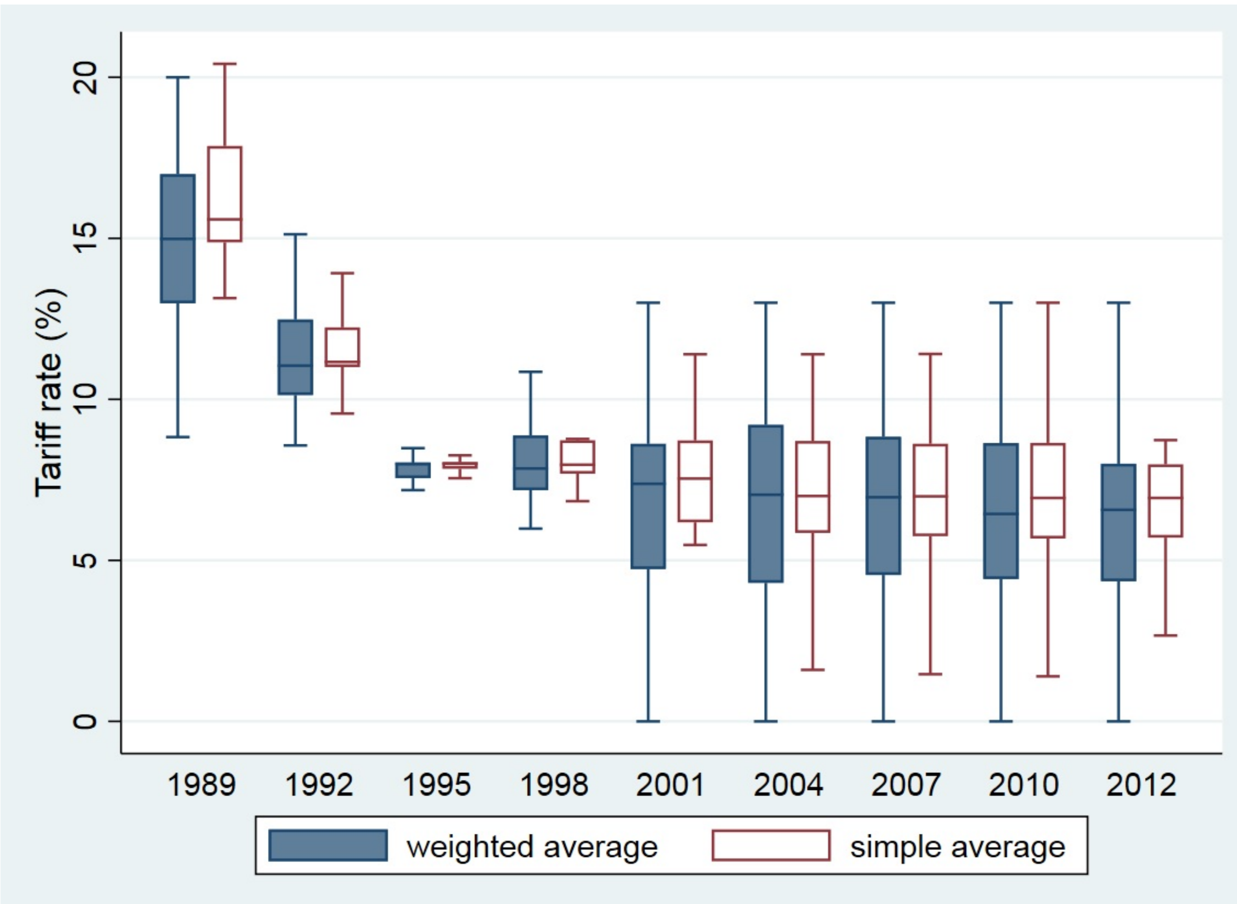
Notes: This figure presents a time trend in the logarithm of average consumer price index with a base year 2010.

Figure 6: Time trend in Average Tariff Rates



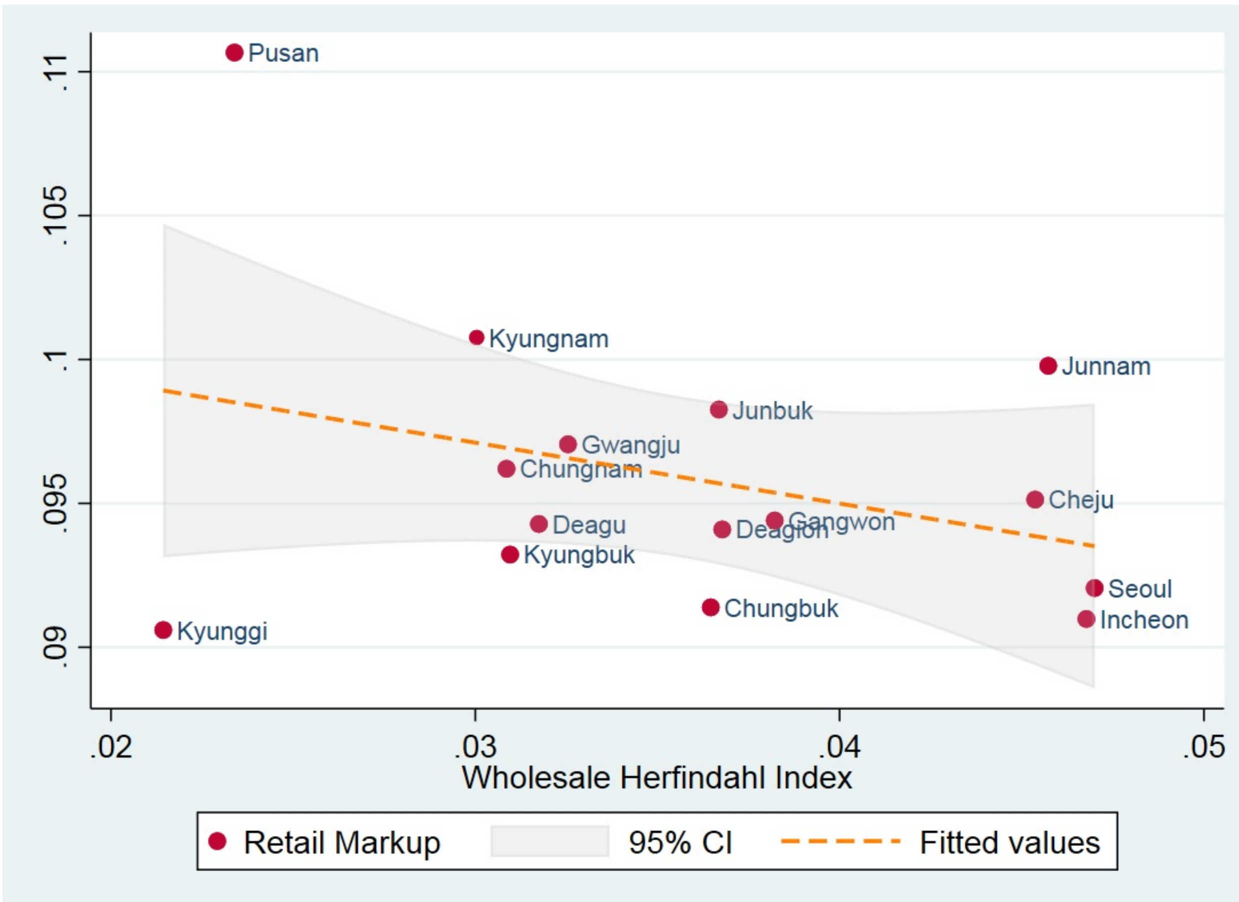
Notes: This figure illustrates the evolutions of Korea's average tariff rates from 1998 to 2012. *Simple average* shows a trajectory of mean values of sectoral (simple) average tariff rates. *Weighted average* describes a trajectory of mean values of sectoral weighted averages. Weights used to measure sectoral weighted averages are one-year lagged import shares in total values.

Figure 7: Tariff Dispersions across Sectors



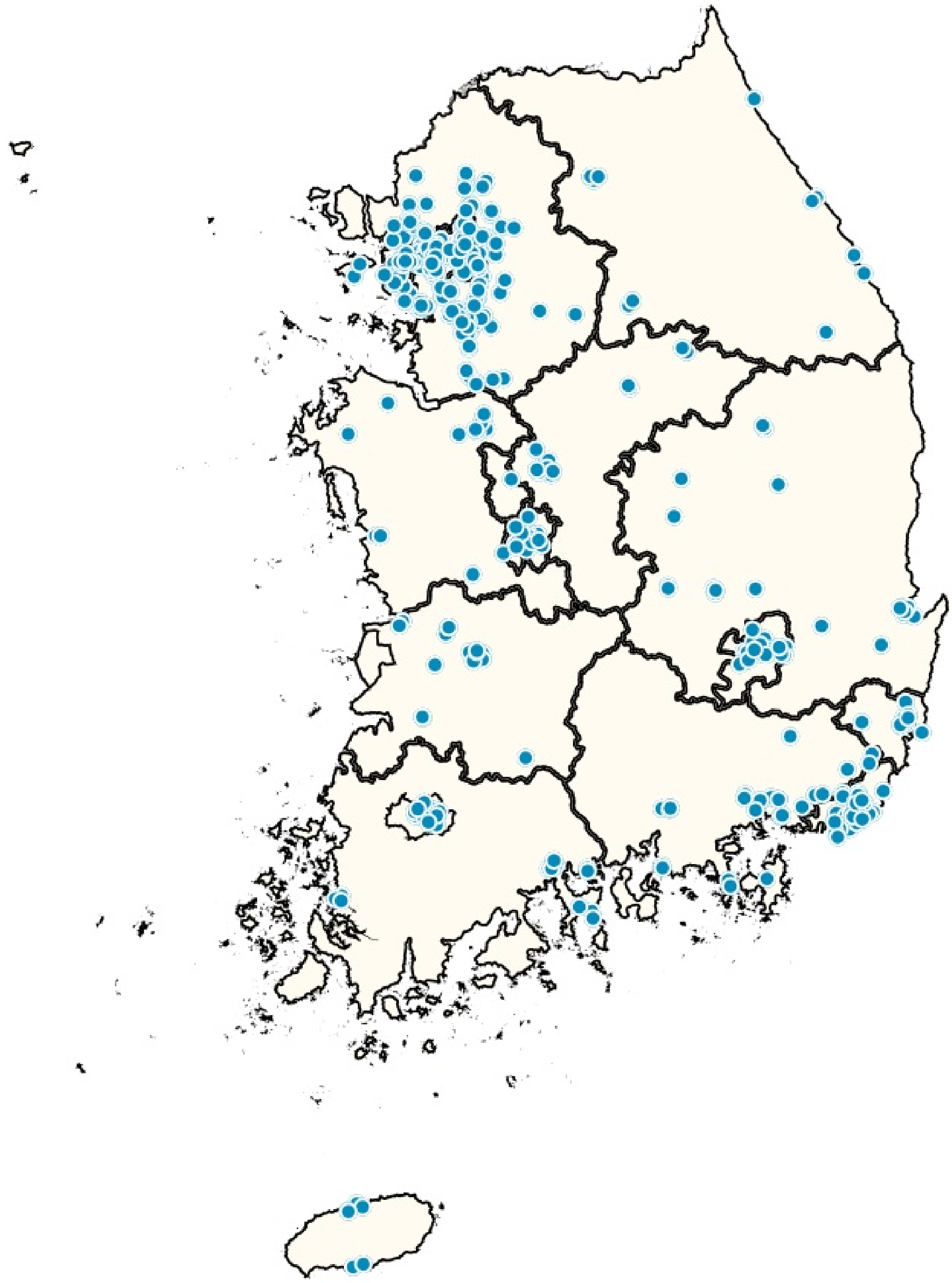
Notes: This figure describes how sectoral dispersions in tariff rates evolve during the period of 1989 to 2012.

Figure 8: Association Between Retail Markups and Wholesale Herfindahl Index



Notes: This figure shows the correlation between the average retail markup and the average Herfindahl index of the wholesale sector across years. It demonstrates that the two variables are negatively associated. With regard to the Wholesale Herfindahl index, I use the empirical index obtained by regressing the analytic Herfindahl index on province- and year fixed effects. It aims both to account for heterogeneities and to reflect the modified index which eventually fits for the first stage IV regression model including the fixed effect terms.

Figure 9: Large Discount Stores in Korea (2010)



Notes: This figure shows the spatial distribution of large discount stores in Korea. Large discount stores are a fairly new retail store format that has proliferated since the mid-1990s in Korea. The spatial distribution of the particular store format illustrates well how locations of retail stores relate to populations.

Appendix Tables

Table A1. Retail Markup Channel, *First Stage IV results*

Dependent Variable:	*Retail Markup* (1)	*Retail Markup* $\times$ *Tariff* (2)
Population_{t-10}	0.000257	0.00430
	(0.000398)	(0.00293)
Wholesale Herhindahl Index	-0.0284***	0.141***
	(0.00486)	(0.038)
$\text{Population}_{t-10} \times$ Tariff	0.00000718	-0.000244***
	(0.0000108)	(0.0000931)
Wholesale Herhindahl Index $\times$ Tariff	-0.000269	-0.0518***
	(0.000542)	(0.00411)
N	4,800	4,800
R^2	0.670	0.972
Clusters	300	300
F-statistics	267.25	239.08

Notes: This table presents estimation results of the first stage IV regression of Equation (12). As mentioned in the main text, there are two endogenous variables: retail markup (μ_{rt}) and retail markup $\times$ Tariff ($\mu_{rt} \times \tau_{it}$). Each column shows the result with respect to each endogenous variable. Population_{t-10} is ten-year lagged provincial populations. Although I do not report here, the fist stage IV regressions obviously include sector-year fixed effect(ϕ_{it}) and province-year specific observed controls ((log) population density and (log) gross regional domestic product). * p-value < 0.10, ** p-value < 0.05, *** p-value < 0.01. Standard errors in parentheses are clustered at the sector-by-year level.

Table A2. Retail Markup Channel, *Analysed by Census Data*

	Weighted Average PCM + Controlling for Town Heterogeneity			
	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
Retail Markup	0.166	6.133*		
	(0.192)	(3.503)		
Retail Markup × Tariff	-0.00090	-0.113*	-0.0052	-0.0892*
	(0.0051)	(0.0680)	(0.0060)	(0.0478)
Province-Year Observables	Yes	Yes	No	No
Sector-Year FE	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	No	No
Province-Year FE	No	No	Yes	Yes
N	1,590	1,590	1,590	1,590
R^2	0.994	0.994	0.994	0.994
Clusters	101	101	101	101
Kleibergen-Paap rk Wald F		3.416		3012.51

Notes: This table suggests that retail market powers decrease tariff pass-through when I use retail markup measure constructed by census data. The measure is weighted average price-cost margins, and I construct them with *town* level market heterogeneity controlled for. More specifically, I apply the same method to establish $\mu_{rt}^{\text{Census}2}$ as described in section 6.1, but the only difference is that I use town fixed effect to account for town-level heterogeneity. Column (1) and (3) are OLS results without IVs, whereas column (2) and (4) show estimation results of (the second stage) IV regressions. Kleibergen-Paap rk Wald F statistics presents the results of week IV test in the case where more than two IVs and clustered standard errors are exploited. While in column (4) the statistics exceed critical values proposed by Stock and Yogo (2005), in column (2) the F statistics do not pass. Thus, unlike column (4), the IV results reported in column (2) have potentially the week instrument variable problem. Province-year observables are (log) population density and (log) gross regional domestic product. * p-value < 0.10, ** p-value < 0.05, *** p-value < 0.01. Standard errors in parentheses are clustered at the sector-by-year level.