Quality Heterogeneity and Misallocation: The Welfare Benefits of Raising your Standards*

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December 2018

Abstract

Domestic regulations on product standards are a key point of contention in modern trade agreements, but frameworks generally ignore their heterogeneous impact across domestic firms. Using data from Chile, we find that more restrictive domestic standards are associated with a reallocation of domestic sales from low- to high-quality firms. Guided by this evidence, we study the welfare effects of quality standards in a model with monopolistically competitive, heterogeneous firms, and a general demand system. Raising the quality standard forces low-quality firms to exit. Welfare improves if low-quality firms over-produce in the market allocation relative to the efficient allocation. A reduction in trade costs reduces the optimal restrictiveness of a standard in a cooperative equilibrium across countries. We estimate our model across Chilean industries and find that in several instances the imposed standard is too restrictive relative to a theoretical upper bound. There is suggestive evidence that a rise in openness has reduced restrictiveness, which supports the efforts to improve agreements on product standards along with traditional trade policy tools.

Keywords: Allocative Efficiency, Minimum Quality Standards, Variable Markups, Trade Policy.

JEL Code: F12, F13, L11.

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*We thank Thibault Fally, Robert Feenstra, Amit Khandelwal, James Lake, Yasusada Murata, Andres Rodriguez-Clare, Katheryn Russ, and Vladimir Tyazhelnikov for suggestions and feedback. We also thank participants at UC Berkeley, UC Davis, UC Santa Cruz, Midwest International Economics Fall Meeting 2017, Aarhus-Kiel Workshop 2017, Danish International Economic Workshop 2018, IOSE 2018, and Moscow International Economics Workshop.

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1 Introduction

Regulations on goods' characteristics are an important tool applied by policy makers. For example, standardization of technical requirements for products is a major priority of the European Commission growth initiative[1]. Governments choose to impose standards in the domestic economy for many legitimate reasons, e.g. standards on auto emissions to counter the negative externality of pollution, or standards in the food industry to protect consumers from disease. Although these regulations are targeted at protecting domestic constituents, product standards are hotly debated in the context of international trade (Maskus et al., 2000; Baldwin et al., 2000; Rodrik, 2018)[2]. Domestic regulations are often viewed as a barrier on market access of foreign firms and for that reason are treated strictly as protectionist (Baldwin et al., 2000; Chen and Novy, 2011; Fontagné et al., 2015).

However, non-discriminatory regulations affect all firms selling to an economy, regardless of their origin, and are aimed specifically at low-quality firms, which are not able to satisfy the requirements imposed by the government. Regulations therefore generate a reallocation of production from low-quality firms to high-quality firms, the welfare implications of which are ambiguous. In this paper, we show that product standards can improve welfare in the presence of firm heterogeneity by reducing the distortions that arise in allocatively inefficient markets. Such distortions originate from the interaction between consumers' preferences and firms' variable market power. Furthermore, we study the effects of trade openness on the optimal degree of restrictiveness of standards. From a policy perspective, we therefore provide a new framework to approach product standards in the negotiation of trade agreements.

We motivate our theoretical framework by documenting the effects of regulations on domestic firms. We compare the sales and survival "premium" that high-quality firms receive, relative to low-quality ones, when industries become more restrictive in their technical requirements. The specification controls for industry-year shocks and time-invariant firm specific characteristics, which is possible with a panel data of Chilean firms and the TRAINS database on domestic regulations on product standards. The main result is that the differences in terms of sales and survival premium of high-quality firms, relative to low-quality firms, are magnified in industries with a larger number of regulations. We interpret this as a reallocation of production from low-quality firms towards high-quality firms.

To study the welfare consequences of such a reallocation, we incorporate regulations on

[1] https://ec.europa.eu/growth/single-market/european-standards_en

[2] In the international trade context, these regulations are labeled technical measures to trade and consist of sanitary and phythosanitary standards and technical barriers to trade (UNCTAD, 2012). The relevance of technical measures in today's trade agreements has surged partly because of the secular decline in tariffs and in the uses of traditional trade policy tools as quotas.

product standards into a standard framework of perfect information, monopolistic competition, and firms that are heterogeneous in quality. We represent the imposition of regulation as a minimum quality standard that a government allows in the market. However, our results generalize to all vertical norms that require the payment of a fixed cost of compliance[3], and do not discriminate between domestic and foreign firms. Consistent with the evidence, raising the standard or, equivalently, making regulations more stringent, forces low-quality firms to exit and, thus, it reallocates production from low-quality to high-quality firms.

There are two opposing effects on welfare. First, the quality standard reduces the total number of varieties available for consumption, as the low-quality varieties exit. In models featuring love for variety (Krugman, 1980), as ours, fewer varieties reduce welfare in conjunction with a weaker competitive environment. Second, the standard improves the average quality in a market, as production moves from low-quality to high-quality firms. The latter channel works to improve welfare when the initial allocation is inefficient.

To provide a general framework to analyze allocative inefficiency, we choose the "Generalized Translated Power" (GTP) preferences proposed by Bertoletti and Etro (2018), which nest the most common preferences used in the trade literature: indirectly additive (IA), directly additive (DA), and homothetic[4]. Given the generality of the demand system, it is striking that the model predicts a non-monotone, hump-shaped relationship between the quality standard and welfare for all parametric specifications. At low levels of the quality standard, improving allocative efficiency dominates the welfare loss from diminishing variety and competition. Eventually the standard becomes too restrictive — above its optimal level — when the welfare loss from diminishing the number of firms offsets the welfare enhancing components of the standard. The optimal level of the standard is determined by parametric assumptions on preferences.

We clarify the mechanisms through which the standard reduces distortions by comparing the market allocation to the socially optimal allocation. Generally, there are three margins through which the market is inefficient: the selection of firms, the quantity produced by each firm, and the number of firms that attempt to enter the market. We limit the analysis to the allocation of production among entrants (the first two margins) by making an assumption common to the literature with firm heterogeneity, that firms draw their quality from a Pareto distribution (Chaney, 2008; Arkolakis et al., 2012, 2017).

[3]We focus on vertical norms, which are easily characterized as being more or less stringent, such as limits on car emissions or on residue levels of pesticides. We abstract from costs associated with the enforcement of the standard that are paid by the government. Moreover, we ignore horizontal norms, which arise when the local firms' differentiated good is adopted as a norm, as electric plugs (Baldwin et al., 2000).

[4]We derive our main theoretical results for a closed economy to abstract from protectionist motives that arise when governments use trade policy to internalize some form of market power (Alvarez and Lucas, 2007).

The distortion reduced by a quality standard is known as "business stealing bias", where too many low-quality firms are active in a market, relative to an optimal allocation[5]. In addition, due to the markup distribution, high-quality firms under-produce and low-quality firms over-produce, relative to an efficient allocation. A necessary condition for such a misallocation is that firms charge variable markups – consumers are willing to purchase low-quality goods provided that those markups are low enough in the laissez faire economy. The standard causes a reallocation of production from low- to high-quality firms, which we label the *composition effect* of the standard.

A consequence of the composition effect is an increase in the weighted average markup due to the reallocation, which improves welfare. However, as the standard reduces the number of competitors, it can generate *anti-competitive* effects, whereby welfare is reduced as surviving high-quality firms increase their markups in response to lower competition. The size of the anti-competitive effects depends on the elasticity of a firm's markup with respect to the number of competitors. The three preferences included in GTP differ in the extent of anti-competitive effects, which are absent in the IA case and are the largest under homothetic preferences. Hence, the model predicts the most restrictive optimal standard under IA, intermediate under DA, and the smallest under homothetic preferences.

A further contribution of the paper is to study the effects of trade openness on the optimal degree of restrictiveness of standards. We abstract from a political economy problem (Bagwell and Staiger, 2001) and consider the optimal standard chosen by two symmetric countries in a cooperative equilibrium. Such a scenario could represent the negotiating phase in trade agreements. The model predicts that trade openness and quality standards are complements: lower trade costs reduce the optimal level of the quality standard. The intuitive mechanism is that trade reallocates production from low-quality non-exporters to high-quality exporters, in a sense doing the work of the standard, and thus reducing its welfare improving ability[6]. Although there is always a low enough quality standard that is welfare improving, as economies open to trade the optimal level decreases. For this reason, this paper offer support of a dual approach for policymakers: pushing towards lower trade costs while lowering unnecessary restrictiveness of quality standards. The result provides a

[5]This intuition is present in Mankiw and Whinston (1986) and Dhingra and Morrow (2016). The business stealing bias dominates another distortion commonly labeled "lack of appropriability", which generates *too little* production from low-quality firms, and occurs when firms cannot fully seize or appropriate the gains from a new variety. Quantitative evidence for this type of misallocation has been highlighted in the aggregate productivity literature (Basu and Fernald, 2002; Restuccia and Rogerson, 2008; Hsieh and Klenow, 2009) and in the trade literature (Edmond et al., 2015; Weinberger, 2017).

[6]The result arises *even* keeping the selection of domestic firms constant, as with IA preferences. We focus only on the quality standard, so that the optimal tariff results of Felbermayr et al. (2013) and Demidova and Rodriguez-Clare (2009) are beyond the scope of this paper.

theoretical justification for the continuous efforts from the WTO of improving the Technical Barriers to Trade Agreement, which has now reached the Eighth Triennial Review[7].

To quantify the effects of product standards, we estimate our model across 40 Chilean manufacturing industries[8]. We note that, although a standard allows for an intuitive theoretical mechanism through which low-quality firms exit, in reality there can be numerous policies that generate the same distributional effect on production. We find a significant presence of measures across Chilean industries. For example, in 2000, the presence of regulations reduced the survival probability of a firm by 40% on average. The restrictiveness of regulations is also heterogeneous across industries: Chemicals, Motor Vehicles, Food, and Books/Journals are consistently the most regulated industries, while Furniture and Apparel are the least regulated, with a reduction in the survival probability of a firm close to 10%.

We conduct a policy-relevant evaluation by comparing the estimated level of restrictiveness with a theoretical upper bound for the restrictiveness of the standard predicted by our model[9]. In 5 out of 38 industries in 2005, we cannot reject the hypothesis that the estimated standard is different than the theoretical upper bound. Hence, in those five industries, the standards are too restrictive in light of our model. Moreover, the number of industries that are too restrictive has declined since 2000. We postulate that an increase in openness could be a factor in reducing the restrictiveness of technical measures, as Chile experiences a boom in trade after 2000. This is exemplified not only by the observed increase in trade flows, but the passage of important free trade agreements with the United States, EU, and China. There is suggestive evidence that the reduction in restrictiveness is largest in the most open industries. This finding reinforces the paper's support for a dual approach of reducing the restrictiveness of standards along with lower trade costs.

Relationship with the Literature. Our paper relates to a growing literature within the trade, industrial organization, and macro fields, on the aggregate consequences of misallocation of production across heterogeneous firms. The presence of misallocation provides a unique channel through which reallocation impacts welfare (Basu and Fernald, 2002; Hsieh and Klenow, 2009; Dhingra and Morrow, 2016). However, practical policy implications are rarely provided as a way to improve upon the observed allocative inefficiency[10]. In this paper, we explore the case where a policy-maker can set a minimum level of quality that is allowed to sell in a market, and generalize the result to the payment of a fixed cost that achieves

[7] https://www.wto.org/english/tratop_e/tbt_e/tbt_triennial_reviews_e.htm

[8] Our estimates of the implied survival restrictions in Chilean industries is similar to Behrens et al. (2018) who use firm revenues/employment to estimate the distortions present within French industries.

[9] This upper bound is the optimal standard in a closed economy with IA preferences. In fact, trade reduces the need for regulations and IA preferences generate the largest restrictiveness of the optimal standard.

[10] Direct policy implications of misallocation include a tax on entry and complicated schemes of firm taxes/subsidies to get the optimal distribution of sales.

the same allocation. Under a plausible set of conditions – governed by the demand faced by firms – regulatory measures imposed on goods' characteristic can raise welfare through an increase in allocative efficiency. The extension of optimality results in Dhingra and Morrow (2016) and Bertoletti and Etro (2018) to a framework with quality differentiation, is a separate contribution of this paper.

An important contribution is to provide a rationale for technical standards that has not been explored in any of the previous literature. Quality standards could be raised to address negative externalities, such as environmental externalities (Parenti, 2016; Mei, 2017), to reduce oligopolists' market power (Baldwin et al., 2000), or to enhance investments in quality upgrading (Gaigne and Larue, 2016). Other reasons, yet to be explored in the context of international trade are information asymmetries or, more generally, information frictions (Schwartz and Wilde, 1985). Last but not least, technical measures could be used as murky protectionism (Baldwin and Evenett, 2009), as studied by Fischer and Serra (2000) in the context of an international duopoly. This paper acts as a complement to the existing literature on rationales for regulations as it is the first to explore the role of inefficient markets. Since our quality standard generalizes to all vertical norms, it allows for more general policy responses.

The trade literature has traditionally considered regulations on product standards as a form of barriers to trade that primarily impacts the extensive margin of firms. In this vein, studies have relied on export flows to show that exporters, and in particular the smallest ones, from a specific origin (e.g France) are less likely to sell a product to destinations that impose relatively more regulations in those products (Fontagné et al., 2015; Fernandes et al., 2015; Ferro et al., 2015). We separate from this literature and examine the effect on domestic firms instead, with a focus on the distribution of firm sales[11]. Our approach fits with the emphasis on firm selection and reallocation of production that are integral to gains from trade when firms are heterogeneous and compete monopolistically (Melitz, 2003).

This paper is organized as follows. Section 2 presents the stylized facts that motivate our modeling choice. Section 3 describes a framework with generalized translated power preferences and quality differentiation, where a policy maker may impose a quality standard. Section 4 shows the results from estimating the model. Section 5 concludes.

[11]The technical measures we examine are imposed on all firms in the economy, not necessarily intended as a trade restriction.

2 Motivational Evidence

The theory in Section 3 frames regulations on product characteristics as a quality standard. Firms with quality below the standard exit, so that their production is reallocated to the rest of the firms in the industry. The measure of active firms, their average quality, and the distribution of sales across these firms determine the level of consumer welfare. In this section, we aim to motivate this approach with firm-level data that allows us to observe survival and sales distributions at the finest industry disaggregation available (4-digit ISIC). We take a balanced panel of Chilean firms and provide evidence of a relationship between the imposition of regulations at the industry level and a growing differential of survival and sales across low- and high-quality firms. In industries with a higher level of quality standards, there is indeed stronger selection of high-quality firms, and the sales distribution is skewed more heavily towards the high-quality firms.

2.1 Data

Detailed Database of Non-Tariff Measures. In order to map our regulations to the data, we make use of the prevalence of technical measures. Technical measures are *domestic* regulations that the WTO interprets as possible barriers to market access. With the secular decline in import tariffs, trade economists have pointed towards technical measures as an increasingly relevant subject in trade agreements (Maskus et al., 2000; Baldwin et al., 2000). Specifically, sanitary and phytosanitary (SPS) and technical barriers to trade (TBT) are now the most crucial technical policies (UNCTAD, 2017).

TRAINS has recently made available a comprehensive database of technical measures imposed by WTO members[12]. The database includes *all* domestic regulations found in official texts that can be classified as non-tariff measures (NTMs)[13]. The 2012 NTM classification separates measures into 16 chapters (labeled A-P), and we make use of the first two chapters: SPS and TBT, to construct our measure of quality regulation. These two chapters are defined by UNCTAD (2017) as "technical measures," and therefore fit most closely with our quality standard in the theory. These are the regulations that Ferro et al. (2015) and Fontagné et al. (2015) have shown to primarily reduce the extensive margin of exporters. An important advantage of this data is that the technical measures apply to both imported goods and locally-produced goods and, thus, do not discriminate between domestic and foreign firms[14].

[12]The web application to retrieve the data is available at http://i-tip.unctad.org/.

[13]TRAINS collects official measures imposed by countries that might affect international trade, that are mandatory, and are currently applied. National governments or local consultants hired by the World Bank collect regulations from official government sources, such as Customs Agencies or Government Ministries.

[14]In contrast, the Specific Trade Concerns data used in Fontagné et al. (2015) lists concerns raised by

From the reporter (imposing country)-product-NTM code-year data, we construct a measure of industry restrictiveness, labeled TM, that can be merged to our domestic production data. First, we allow each product to have at most one regulation imposed per NTM (2 digit) chapter in each year, and aggregate the number of regulations for each 4-digit ISIC (revision 3) industry. To control for the number of products in each industry, we divide the previous sum by the number of HS6 products in the 4-digit industry[15]. Finally, the industry classification allows us to merge the measures to the firm data which we describe next. Table 4 in Appendix 6.1 lists the top 25 industries ranked by the restrictiveness in the 1995-2007 period. Unsurprisingly, these rankings are populated by food and pesticide products due to the presence of SPS regulations, along with chemicals and equipment machinery.

Chilean Firm Data. The Chilean data is a census of a panel of firms with more than 10 employees from 1995 to 2007, provided by Encuesta Nacional Industrial Anual (ENIA, National Industrial Survey) and collected by the National Institute of Statistics (INE). Each firm is classified with a 4-digit ISIC industry. There are approximately 5,000 firm level observations per year and firms are tracked across time with a unique identification number. The census includes detailed firm data such as total sales, value/usage of its factors, etc.

The key limitation, as faced by previous literature, is the lack of an explicit measure of quality. We rely on input measures to construct a proxy for quality: we take the firms' capital stock, labor costs, and intermediate input costs, and divide each by the number of employees. Higher capital intensity, average wage per worker, and average material input costs all arguably correlate with quality, and have been used in previous studies (Kugler and Verhoogen, 2012)[16]. Hallak and Sivadasan (2013) use the same quality proxies for Chile, however they complement these with Indian product-level data that allows them to also use the adoption of ISO 9000 certification and input/output prices. When they investigate the exporter quality premium conditional on size, they find very similar results across all of the quality proxies. Therefore, the three proxies in our paper are likely capturing very similar attributes as the more direct quality measures, although access to product price data would allow us for better proxies of quality. In order to create a quality indicator, we label a firm as "high-quality" if it is above the median in the quality proxy within its industry in 1995.

trading partners. To safeguard against discriminatory technical measures, in the main specification we drop "Pre-shipment Inspections" (Chapter C in the NTM classification, which is often represented as a technical measure), as well as classifications A1 and B1 within the SPS and TBT chapters respectively, as these are most likely to include regulations that only affect imported goods.

[15]Re-defining TM without controlling for the number of HS6 products does not change the results

[16]The three proxies are positively correlated with a ratio of skilled workers to unskilled workers, using a (rough) measure in the Chilean data that labels a category of workers as unskilled (or "no calificados" in Spanish). Not surprisingly, the correlation is strongest with the wages per worker quality proxy.

2.2 Technical Measures and Chilean Firms

The data described above allows us to test the distributional effects of technical measures within industries. To do so, we run the following specification:

$$y_{fit} = \alpha_{it} + \alpha_f + \beta_M TM_{it} * Quality_f + \beta_X X_{it} * Quality_f + \epsilon_{fit}, \tag{1}$$

where y_{fit} is a performance measure for firm f in industry i at year t which includes log domestic sales and a dummy for positive sales ("survival premium"). TM_{it} is the measure of industry restrictiveness based on the imposition of SPS and TBT measures as reported in Table 4. The main coefficient of interest is β_M, which identifies the high- versus low-quality differential response to the imposition of regulations in an industry-year.

We include industry-year (α_{it}) and firm (α_f) fixed effects to control for the variety of industry and macroeconomic shocks, plus time invariant firm characteristics. This restrictive specification only captures the relative firm outcomes that are due to changes in technical measures and not due to the various industry characteristics that might drive the firm sales distribution[17]. The time-varying controls, $X_{it} * Quality_f$, capture changes in non-regulatory industry characteristics that might drive relative outcomes between high- and low-quality firms. These include an interaction of industry openness with the quality indicator to control for differences in competition introduced by trade, and an interaction of the quality indicator measure with the level of import tariffs at the industry level[18].

In our main specification, we consider a balanced panel of firms. We keep only firms alive in 1995 and construct a balanced panel where a firm is given a survival dummy equal to 0 if it does not sell in that year. This follows the specification in Fontagné et al. (2015) and allows us to interpret the firms in the first year as the "potential" producers. Firms are assigned a quality indicator based on being above or below the median in 1995[19]. To some degree the results on the "survival" outcome are affected by the fact that firms with less than 10 employees are not forced to participate to the survey. However, given that we find exit to be more prevalent among the smallest firms and the sign on relative survival (β_M) is in the direction that we expect, the censoring of the data likely *understates* the magnitude of the firm churning.

[17] Our specification controls for time-varying industry characteristics, as well as time-invariant firm characteristics, that are correlated with the sales distribution. For example, differences in product differentiation and demand elasticities across firms and industries are controlled for with the fixed effects.

[18] As described below, import tariffs declined in this period, although mostly uniformly across industries.

[19] For the specification with sales as an outcome, our results are robust to using the unbalanced panel.

Results. Table 1 reports the main motivational results. We include the controls detailed above but only report the main coefficients of interest β_M (a full table is in the appendix). The results on domestic sales suggest that the ratio of sales between high- and low-quality firms is magnified when industries become more regulated. The coefficient in the first row of column (1) implies that imposing a regulation for every product in an industry results in a 1.5% larger sales difference between an average high-quality firm relative to the average low-quality firm, with a p-value of 0.01. The three proxies for quality yield similar results, although they are the most precise when using capital intensity. The last three columns suggests that the survival of high-quality firms relative to low-quality firms is also higher in more regulated industries. In this case, only the capital intensity results are statistically significant below the 5% level, but the sign is positive and of similar magnitude for all proxies of quality.

Table 1: Firm Sales (top) and Survival (bottom) Heterogeneity - by TMs in Industry

	Log Domestic Sales			Survival		
	(1)	(2)	(3)	(4)	(5)	(6)
TM*Quality (Capital/L)	0.015**			0.012**		
	(0.006)			(0.005)		
TM*Quality (Wage/L)		0.012**			0.009	
		(0.006)			(0.006)	
TM*Quality (InputValue/L)			0.008			0.006
			(0.006)			(0.004)
Fixed Effects	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y
R^2	0.955	0.955	0.955	0.645	0.644	0.643
# Observations	44220	44220	43789	69679	69679	68924

Results from OLS estimation of (1). TM_{it} (restrictiveness) is measured at the 4 digit ISIC industry level. The total number of measures in each industry-year are summed and then divided by the number of HS6 products in the industry. Each row interacts the TM measure with a dummy for quality, where quality is proxied by capital per worker, total wages per worker, and input expenditure per worker respectively. Firms are high-quality if the quality proxy is above the median in that industry in 1995. For the results on survival, all firms alive in 1995 are "potential" producers in all years, which is why the number of observations is much larger. In all specifications we include an interaction of industry openness with the quality indicator, an interaction of the quality indicator measure with the industry import tariff, plus firm and industry-year interacted fixed effects. Standard errors – clustered by 4-digit industry – are in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

The results of this section are consistent with the following intuition: in industries with more regulations there is more exit of low-quality firms, and this reallocates production to the high-quality firms. We do not interpret the results as a causal relationship between the imposition of technical measures and sales/survival heterogeneity, but instead we motivate the model by highlighting the existence of this relationship. These results also line up with the findings in Fontagné et al. (2015) that regulations hurt the small exporters the most, assuming that the small exporters are also viewed as selling lower quality.

In the appendix, we show that there is a strong correlation in the data between all the

quality proxies and firm sales, which allows us to interpret a quality standard as essentially eliminating firms in the left tail of the sales distribution. Furthermore, we also document a strong relationship between all the quality proxies and a measure of TFP. This result is important because our theoretical implications with quality heterogeneity can be translated to productivity heterogeneity but only in the familiar case that quality is proportional to productivity. These relationships are also consistent with Hottman et al. (2016), which find that product "appeal" is the most important component of sales heterogeneity. In summary, the empirical findings motivate our theoretical framework, with firms differentiated by quality, and a standard that eliminates the lowest quality firms.

As there are several caveats to keep in mind, in Section 4, in order to not rely on the possibly flawed data on technical measures, we estimate industry restrictiveness from the structure of the model by backing out the implied restrictiveness given the empirical sales distribution. Next, we briefly describe several robustness results to the above empirical specification that are reported in Appendix 6.1.

Robustness. A common issue with data on regulations is the high level of measurement error. For instance, there could be a mismatch between the date of initial enforcement of a regulation, and the date of its listing in the dataset. To address the concern, we run a specification where regulations are aggregated across all years so that there is *one* restrictiveness measure for each industry. In this case, the specification is a repeated cross-section, with sales as the outcome within industry-year, and ran on an unbalanced panel[20]. We find that more regulated industries exhibit higher skewness in sales towards high-quality firms, which suggests that the timing of when regulations are listed does not drive the results.

The results rely on the implicit assumption that technical measures are non-discriminatory. In fact, regulations must fit this criteria to be legal under WTO rules, and we attempt to omit technical measures that might be more heavily weighted towards importers. To test this assumption, we create our TM variable using only a subset of technical measures dropped from the main analysis that might be aimed at importers – those classified by UNCTAD as "Prohibitions/restrictions of imports for SPS reasons" and "Prohibitions/restrictions of imports for objectives set out in the TBT agreement". Since these are the measures least likely to affect domestic firms, we expect to not find the same type of evidence for reallocation. In fact, we find the opposite result of our baseline specification: in 5 of the 6 interactions the

[20]The main drawback in this case is that we cannot control for firm fixed effects. Since the regulations are aggregated from the HS 6 product level, firms within the same 4-digit ISIC might actually be exposed to different levels of regulation. We add an interaction with industry trade elasticities (from Broda and Weinstein (2006)) to control for the effect of demand characteristics on the sales distribution – which was controlled for in the previous specification by firm fixed effects.

coefficients are negative. We caution however that these results have large standard errors as there are few technical measures that fit this definition[21].

Our dataset contains both vertical and horizontal norms, although our theoretical framework only considers vertical norms. As a way to deal with this issue, we replace the set of technical measures used in specification (1) with more specific subsets of technical measures (Table 9 in the appendix). For example, we construct two versions of TM using SPS and TBT separately. Either type of technical measure is associated with a widening of sales dispersion and lower survival for low-quality firms. However, the results are stronger in the case of SPS regulations – mostly due to much smaller standard errors. This could be due to the fact that SPS measures are more prevalent than TBT ones, which we find to be the case. However, this suggests the results are likely not driven by horizontal norms, which are arguably more likely to come in technical regulations under TBT[22].

Finally, we include an IV specification in the appendix where TM is instrumented using the TM measure in *Peru* interacted with the same quality indicator. This specification minimizes the concern that Chilean regulations may reflect Chilean consumers preference for quality[23]. Although the results are weaker, the qualitative story stays the same.

3 Theory

This section builds a theory for the welfare effects of regulations on good's standards. We begin by presenting the description of the environment, with a standard supply side and a general demand system that nests several preferences common in the trade literature. Then, we proceed by allowing a policy maker the option of imposing a quality standard, whose effects on the distribution of firms are consistent with the evidence documented in the previous section. We derive an expression for welfare as a function of the standard, and find that a standard more restrictive than the market allocation is always optimal. We discuss the sources of market distortions that a minimum quality standard reduces, and identify the features of each type of preferences that cause shifts in the magnitude of the optimal standard in an economy. Given the generality of this demand system, our welfare results provide a strong motivation for the rationale of a quality standard. We end the section with a study of the effects of trade openness on the optimal quality standard that two countries

[21] One coefficient is positive and large (though insignificant), but overall the results do not point to the same reallocation effects present with the other measures.

[22] The presence of measures based on horizontal norms likely biases our results towards zero. These measures do not discriminate on any attributes related to quality, which means that "treated" industries will receive no distributional impact.

[23] We use the TM in Peru because we find this country to be closest to Chile in terms of regulatory structure across industries (and therefore the F-stat in the first stage is very large).

cooperatively choose.

3.1 Framework

Consider a closed economy, where L consumers enjoy the consumption of varieties of a differentiated good. We normalize per capita income to 1. The varieties are produced by a mass of single-product firms, which differ in terms of their quality z. We assume that quality z is a demand shifter: consumers exhibit a higher willingness to pay for higher quality goods. There is perfect information: consumers, firms, and the government costlessly distinguish between the quality offered in the market[24].

As in the Melitz (2003) model, there is a pool of potential entrants. Upon entry, firms pay a fixed cost of entry f_E in labor units and discover their quality z. Quality is drawn from an unbounded Pareto distribution with shape parameter κ and shift parameter b. The CDF of the distribution is $H(z) = 1 - \left(\frac{b}{z}\right)^{\kappa}$, while the pdf is $h(z) = \frac{\kappa b^{\kappa}}{z^{\kappa+1}}$. Only a mass J of firms pays the fixed cost of entry. Free entry drives expected profits equal to f_E.

The market is monopolistically competitive. All firms produce their goods with the same marginal cost of production c, in labor units. These assumptions imply that size heterogeneity is linked to the exogenous quality draws. The direct mapping of quality to size might seem stark, but it is a convenient feature that is also present in Kugler and Verhoogen (2012) and finds quantitative support in the empirical findings of Hottman et al. (2016).

3.2 Consumer and Firm Problems

3.2.1 Consumer's Problem

We adopt the Generalized Translated Power (GTP) preferences proposed by Bertoletti and Etro (2018):

$$U = \int_{\Omega} \left(az(\omega)\xi q(\omega)) - \frac{(\xi q(\omega))^{1+\frac{1}{\gamma}}}{1+\frac{1}{\gamma}} \right) d\omega + \frac{\xi^{-\eta} - 1}{\eta} \tag{2}$$

where $a > 0$ and $\gamma \geq 0$ are constants, $q(\omega)$ is the quantity consumed of variety ω, $z(\omega)$ is a variety specific demand shifter, which we interpret as quality, and Ω is the set of varieties available for consumption. ξ is a quantity aggregator that is implicitly defined as:

$$\xi^{-\eta} = \int \left(az(\omega)\xi q(\omega) - (\xi q(\omega))^{1+\frac{1}{\gamma}} \right) d\omega \tag{3}$$

[24]Papers on regulations often introduce an *ad hoc* externality, usually pollution or public health, to justify the imposition of a standard (Parenti, 2016; Mei, 2017). In this spirit, we can also think of our quality measure as a proxy for how healthy a product is. Despite consumers knowing and appreciating the healthiness of a good, our model features a market allocation that generates over-production of unhealthy products.

The GTP utility follows the generalized Gorman-Pollak demand system[25], and nests several preferences based on the value of the parameter $\eta \in [-1, \infty]$. For $\eta = -1$, preferences are indirectly additive (IA) as described by (Bertoletti et al., 2018). For $\eta = 0$, preferences become homothetic with a single aggregator. For $\eta \to \infty$, preferences become directly additive (DA), and generalize the preferences used by Melitz and Ottaviano (2008)[26]. Fally (2018) describes the regularity conditions for these preferences[27].

The consumer's budget constraint is:

$$\int_{\Omega} p(\omega)q(\omega)dz \leq 1$$

where $p(\omega)$ is the price of variety ω and per capita income is normalized to 1. The consumer chooses $q(\omega)$, $\omega \in \Omega$, to maximize its utility subject to the budget constraint. Consumer's inverse demand is:

$$p(\omega) = \xi^{1+\eta}\left[az(\omega) - (\xi q(\omega))^{\frac{1}{\gamma}}\right] \tag{4}$$

3.2.2 Firms' Problem

Given the quality draw z, a firm maximizes its profits by choosing quantity $q(z)$ taking ξ as given. Profits are given by:

$$\pi(z) = L\xi^{1+\eta}\left[azq(z) - \xi^{\frac{1}{\gamma}}(q(z))^{1+\frac{1}{\gamma}}\right] - Lcq(z) \tag{5}$$

The first order condition with respect to $q(\omega)$ equals:

$$\xi^{1+\eta}\left[az - \left(1 + \frac{1}{\gamma}\right)(\xi q(z))^{\frac{1}{\gamma}}\right] = c$$

and setting $q(z^*) = 0$ yields the market determined quality cutoff:

$$z^* = \frac{c}{a}\xi^{-(1+\eta)} \tag{6}$$

For a quality level below the cutoff $z < z^*$, a firm has zero demand. The relationship between the cutoff and η will be key in comparing our results across the types of preferences because the demand faced by each firm is governed by the firms' quality relative to the market cutoff. The quality cutoff in the IA case ($\eta = -1$) only depends on income (normalized

[25] Gorman (1972), Pollack (1972).

[26] The case where $\gamma = 1$ generates linear demand as in the separable case of Melitz and Ottaviano (2008).

[27] In Web Appendices we make available interesting non-GTP cases such as constant elasticity demand and separable variable elasticity demand (e.g. Simonovska (2015)).

to one): $z^*_{\text{IA}} = \frac{c}{a}$. The cutoff for homothetic preferences ($\eta = 0$) depends only on the number of competitors and is independent of income: $z^*_{\text{H}} = \frac{c}{a}\xi^{-1}$. In the DA case, the market determined cutoff depends on *both* income and the number of competitors. Given the relationship between ξ and the marginal utility of income λ[28], for $\eta \to \infty$, $z^*_{\text{DA}} = \frac{\lambda c}{a}$.

Substituting the cutoff (6) into the first order condition yields the optimal quantity:

$$q(z) = \left(\frac{a\gamma}{1+\gamma}\right)^{\gamma} \frac{(z^*)^{\gamma}}{\xi} \left(\frac{z}{z^*} - 1\right)^{\gamma} \tag{7}$$

As $q(z)$ is increasing in z, active firms with higher quality sell larger quantities of their products. Substituting (7) into (4) yields the optimal pricing rule:

$$p(z) = c \underbrace{\frac{1}{1+\gamma}\left(\frac{z}{z^*} + \gamma\right)}_{\text{Markup}} \tag{8}$$

Markups are increasing in z: higher quality firms charge higher markups. Such prediction receives empirical support from Bastos and Silva (2010), Martin (2012), Dingel (2015), and Manova and Zhang (2017).

Firm z revenues $r(z)$ and profits $\pi(z)$ are given by:

$$r(z) = \frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma} \frac{(z^*)^{\gamma}}{\xi} \left(\frac{z}{z^*} - 1\right)^{\gamma} \left(\frac{z}{z^*} + \gamma\right) \tag{9}$$

$$\pi(z) = \frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma} \frac{(z^*)^{\gamma}}{\xi} \left(\frac{z}{z^*} - 1\right)^{1+\gamma} \tag{10}$$

3.3 Quality Standard and Welfare

The government of the closed economy sets a minimum quality standard $\bar{z} \geq z^*$, such that a firm with quality $z < \bar{z}$ is not allowed to sell in the economy. The quality standard is a vertical norm (Baldwin et al., 2000): $\bar{z}$ can be easily interpreted as more or less restrictive. Since firms' quality is exogenously determined, the policy only affects the selection of firms into the domestic market. In particular, the larger $\bar{z}$ becomes, the more low-quality firms are forced out of the market. The model is consistent with the evidence of Section 2.

Our results generalize to all vertical norms that require the payment of a fixed cost of compliance by all firms. This is an important generalization because it allows for a separate way to impose the standard: a policy-maker can impose a fixed production cost that generates the same exit of low-quality firms as $\bar{z}$. In appendix 6.2.4, we investigate the

[28] $\lambda = \frac{1}{y}\int \left(az\xi q(\omega) - (\xi q(\omega))^{1+\frac{1}{\gamma}}\right) d\omega = \frac{\xi^{-\eta}}{y}$, where λ is the multiplier on the resource constraint.

case where the standard is imposed as a fixed cost, which merely generates a downward shift in the level of the optimal standard. We choose to model $\bar{z}$ as a direct policy tool because, in Section 4, we will be able to estimate its restrictiveness regardless of the level of the fixed cost. We abstract from any costs associated with enforcing the standard by the government, which would be hard to quantify, and would reduce the welfare benefits of standards.

We abstract from firms endogenously choosing their quality à la Feenstra and Romalis (2014) because such an assumption would not generate any additional sources of distortions. Moreover, in the Feenstra and Romalis (2014) framework, raising the standard would raise the quality of the smallest surviving firms, which would raise their revenues and thus reduce the sales difference between high-quality and low-quality firms, contrary to the evidence of Section 2. Finally, assuming that to produce higher quality goods a firm incurs higher marginal costs, as in Manova and Zhang (2017), only quantitatively affect the results (see Appendix 6.2.5).

It is convenient to write our variables as a function of $g = \frac{\bar{z}}{z^*} \in [1, \infty)$, a measure of the restrictiveness of the quality standard. If $g = 1$, the standard is ineffective: the market-determined quality cutoff z^* is equal to the minimum allowed $\bar{z}$. For $g > 1$, the government is enforcing a higher quality standard than the one determined by the market. The measure g is related to the probability of a firm being active under the restriction, relative to the same probability without the restriction: $\frac{P(z \geq \bar{g} | g > 1)}{P(z \geq \bar{g} | g = 1)} = g^{-\kappa}$.

3.3.1 Market Aggregates

We start with the market aggregates necessary to compute welfare. Details on the derivations are relegated to the appendix. The equilibrium quality cutoff z^* can be represented as a function of the restrictiveness of the standard g and parameters:

$$z^* = \left[\frac{L c^{\frac{\eta}{1+\eta}} \gamma^{\gamma} b^{\kappa} a^{\gamma + \frac{1}{1+\eta}}}{f_E (1+\gamma)^{1+\gamma}} g^{-\kappa} G_1(g) \right]^{\frac{1}{\kappa - \gamma - \frac{1}{1+\eta}}} \tag{11}$$

where $g^{-\kappa} G_1(g)$ is decreasing in g[29]. The parameter η controls the elasticity of the quality cutoff with respect to market size L and marginal costs c. In particular, the elasticity of the cutoff with respect to size is $\frac{\partial \ln z^*}{\partial \ln L} = \frac{1}{\kappa - \gamma - \frac{1}{1+\eta}}$. An increase in market size induces selection effects, namely it increases the minimum level of quality allowed by the market, if such an elasticity is positive. Such a condition is satisfied for homothetic ($\eta = 0$) and DA preferences

[29] $G_1(g) = \kappa g^{1+\gamma} \left[\frac{F_1(g)}{\kappa - \gamma - 1} - g^{-1} \frac{F_2(g)}{\kappa - \gamma} \right]$. $F_1(g)$ and $F_2(g)$ are two hypergeometric functions given by $F_1(g) = {}_2F_1 \left[\kappa - \gamma - 1, -\gamma; \kappa - \gamma, g^{-1} \right]$ and $F_2(g) = {}_2F_1 \left[\kappa - \gamma, -\gamma; \kappa - \gamma + 1, g^{-1} \right]$. We restrict the parameter space such that $\kappa - \gamma - 1 > 0$.

($\eta = \infty$). However, under IA preferences ($\eta = -1$), where the cutoff is only dependent on income, there are no selection effects due to market size[30].

Substituting (11) into (6) yields the aggregator ξ:

$$\xi = \left[\frac{Lb^{\kappa}a^{\kappa}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}g^{-\kappa}G_1(g)\right]^{-\frac{1}{(1+\eta)(\kappa-\gamma)-1}} \tag{12}$$

which equals one under DA preferences, decreases in g under IA preferences, and increases in g under homothetic preferences. The aggregator ξ is a quantity shifter that affects the volumes of production, along with z^*, of all surviving firms. Hence, the quality standard has a partial negative effect on the volumes produced under IA preferences, and a partial positive effect under homothetic preferences.

Finally, the mass of entrants J is independent of η:

$$J = \frac{L}{f_E}\frac{G_1(g)}{G_2(g)} \tag{13}$$

and is increasing in the restrictiveness of the standard. As an increase in g increases the average profits in the economy, more firms enter. However, the total number of active firms in the economy $N = P(z > \bar{z})J$ is declining in the restrictiveness of the standard.

3.3.2 Welfare

We are now ready to express welfare as a function of the quality standard. After integrating over the two terms in (2) (see appendix), the utility becomes:

$$U = \frac{az^*\xi}{c}\left[(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)} + \frac{1}{\eta}\right] - \frac{1}{\eta} \tag{14}$$

The term $(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}$, which is a component of the average utility, is always increasing in g[31]. On the other hand, the product of the quality cutoff z^* and the aggregator ξ is declining in g. Using the cutoff condition (6) and the equilibrium value of ξ (12) yields the utility of consumers as a function of g:

$$U = \left[\frac{Lb^{\kappa}a^{\kappa}\gamma^{\gamma}g^{-\kappa}G_1(g)}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}\right]^{\frac{\eta}{(1+\eta)(\kappa-\gamma)-1}}\left[(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)} + \frac{1}{\eta}\right] - \frac{1}{\eta} \tag{15}$$

[30] The elasticity of the cutoff with respect to marginal costs c, similar to the income effects of Bertoletti and Etro (2018), is $\frac{\partial \ln z^*}{\partial \ln c} = \frac{\eta}{(1+\eta)(\kappa-\gamma)-1}$. $\frac{\partial \ln z^*}{\partial \ln c} = 0$ for homothetic preferences, as in Melitz (2003), $\frac{\partial \ln z^*}{\partial \ln c} = 1$ for IA preferences ($\eta = -1$), and $\frac{\partial \ln z^*}{\partial \ln c} = \frac{1}{\kappa-\gamma} < 1$ for DA preferences.

[31] $G_2(g) = \kappa g^{1+\gamma}\left[\frac{F_1(g)}{\kappa-\gamma-1} + \gamma g^{-1}\frac{F_2(g)}{\kappa-\gamma}\right]$ and $G_3(g) = \kappa g^{1+\gamma}\left[\frac{F_1(g)}{\kappa-\gamma-1}\right]$.

As shown in Figure 1, a minimum quality standard can improve welfare across the three preferences nested into GTP. In fact, the relationship between welfare and the standard is hump-shaped. There are two opposing welfare effects that generate such a relationship. First, in the presence of a quality standard, the selection of firms is determined by the government imposed $\bar{z}$, and not by the market cutoff z^*. We call this the *composition effect* of the standard: regardless of the preferences, the exit of low-quality firms reallocates production towards the surviving high-quality firms. Such a reallocation is welfare improving.

Second, the quality standard reduces the number of varieties available for consumption, which is welfare reducing under the assumption of love for variety. Furthermore, the reduction in the number of varieties may cause a change in the markups of surviving firms, through a change in z^* (8). The effects of the standard on markups of surviving firms depend on the preferences used and, in particular, on the elasticity of the market cutoff with respect to L. Under IA preferences, in which such an elasticity is zero, the standard leaves the markups of surviving firms unchanged. Under DA and homothetic preferences, the standard increases the markups of surviving firms. We call this the *anti-competitive* effect of the standard: it operates under homothetic and DA preferences, but is silent under IA preferences. Finally, the increase in markups, or anti-competitive effect, is largest under homothetic preferences[32].

For "small" levels of restrictiveness the composition effect dominates the reduction in the number of varieties and the anti-competitive effects. Increasing the standard over its optimal value causes the variety reduction to dominate, and welfare starts falling.

The optimal level of the measure of the restrictiveness of the standard $g^{opt}(\kappa, \gamma, \eta)$ only depends on the parameters κ, γ, and η. The optimal level of the standard $\bar{z}^{opt}$ is then proportional to the market-determined cutoff:

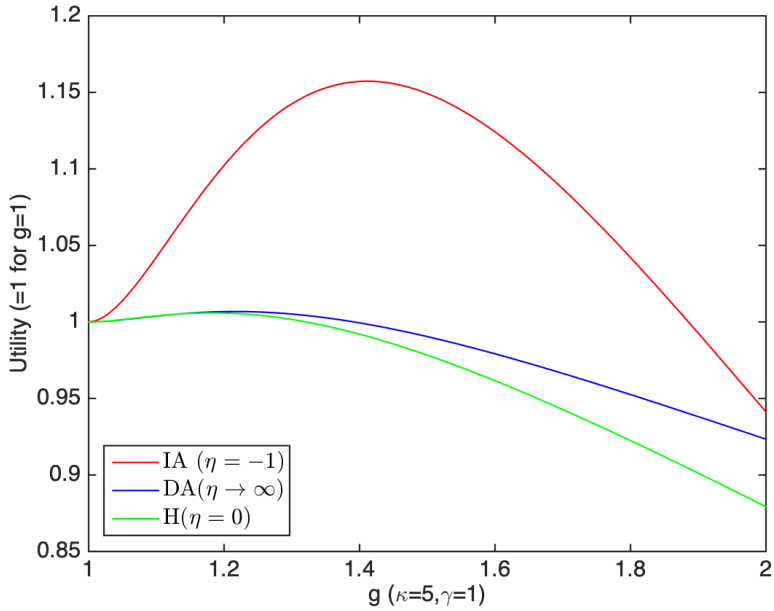
$$\bar{z}^{opt} = g^{opt}(\kappa, \gamma, \eta) z^* \tag{16}$$

If we interpret z^* as a market determined preference for quality, markets with higher preference for quality have higher optimal quality standards while markets with a lower preference for quality have a lower optimal level of $\bar{z}$. To derive some quantitative intuition for the result, let us focus on the IA case, in which z^* is a constant. For $\kappa = 5$ and $\gamma = 1$, welfare is maximized at $g = 1.41$: the government sets a standard which reduces the probability of a firm being active by $|1.41^{-5} - 1| = 82\%$ relative to the market allocation[33].

[32] In fact, one can relate this result to Arkolakis et al. (2017), who show that the effect of *trade costs* on the choke price can be ranked across the same types of preferences.

[33] Under homothetic and DA preferences, z^* is a function of $\bar{z}$. Hence, the reduction in the probability of being active becomes $|(\tilde{g})^{-\kappa} - 1|$, where $\tilde{g} = \frac{\bar{z}}{z^*(1)} = g \left[\frac{g^{-\kappa} G_1(g)}{G_1(1)} \right]^{\frac{1}{\kappa - \gamma - \frac{1}{1+\eta}}}$.

Figure 1: Minimum Quality Standard and Welfare



3.4 Discussion

The most direct way to interpret how the quality standard alters welfare is through the two channels described above. A higher standard lowers the number of varieties available, which lowers welfare, but it raises allocative efficiency. The latter channel raises the measure of average markups in the market allocation closer to the average social markups. That misallocation is reduced when average markups increase might seem counter-intuitive, but in fact allocative efficiency increases as market share is reallocated away from low-quality firms and to high-quality firms, a channel highlighted with productivity heterogeneity in recent work by Edmond et al. (2018), Baqaee and Farhi (2017), and Weinberger (2017).

Since a quality standard $\bar{z}$ can improve welfare because the market allocation is inefficient, the rest of the subsection is devoted to understand in detail which distortions are reduced by a standard, and how these differ across the types of preferences. There are three possible margins through which the market equilibrium is inefficient: entry, selection, and the distribution of markups across active firms. However, the assumption of Pareto distributed quality and monopolistic competition constraints the margins of inefficiency present in our model to the allocation of production across entrants (the latter two)[34].

To understand these two margins, we recap the two biases identified in Dhingra and Morrow (2016) (DM). The first type of distortion is due to lack of appropriability: in making

[34]The mass of entrants J is always efficient in the market allocation (Arkolakis et al., 2017).

their production decision, firms do not take into account the social gains from an increase in variety. Letting z_P^* denote the optimal cutoff chosen by a planner, this "appropriation bias" causes an excessive degree of firms' selection, whereby $z^* > z_{\mathrm{P}}^*$ all else equal. Firm heterogeneity in market power generates the second distortion: in making their production decision, firms do not take into account how their choice alters production and prices of other firms. This "business stealing" effect (DM and Mankiw and Whinston (1986)) reduces selection below the optimum, i.e. $z^* < z_{\mathrm{P}}^*$, because it allows low-quality firms to steal business from high-quality firms[35]. Moreover, the business stealing bias distorts the quantity of production across firms. High-quality firms under-produce as their markups are too high and low-quality firms over-produce, relative to the efficient allocation.

The quality standard affects welfare in two opposing directions in (15), both of which can be understood through its effect on markups. First, the standard raises the *average markup* through a composition effect which works purely through a reallocation of market shares and bring the economy closer to the socially optimal average markup[36]. The standard eliminates low-quality firms, reducing the distortion that affects selection, and furthermore causes a reallocation of production towards high-quality firms, therefore reducing the distortion on the distribution of quantities produced. These are the two inefficiency margins discussed above. Second, the standard reduces the number of varieties. Such a reduction, which is welfare reducing in and of itself, can lead to anti-competitive results. As discussed above, the standard can reduce competition and thus raises the markup of *each surviving firm*. For small values of g, the composition effect dominates for any η[37].

In the following paragraphs, we describe these market inefficiencies that emerge for each of the three specific cases of GTP preferences. To do so, we compare the main variables of interest between the social planner's allocation and the market's allocation. Details of the planner's allocation are in the Web Appendix.

IA Preferences. Under IA preferences ($\eta = -1$), the market allocation always generates a business stealing bias. The ratio of the planner's quality cutoff relative to the market cutoff is always greater than one: $\left(\frac{z_P^*}{z^*(1)}\right)_{IA} = \frac{\kappa-\gamma}{\kappa-\gamma-1} > 1$. As a result, low-quality firms over-produce and high-quality firms under-produce relative to the planner's allocation. The composition effect of the standard reduces the business stealing bias, by forcing the exit of low-quality firms and increasing the production of surviving firms. Although the anti-competitive effect

[35]The Dhingra and Morrow (2016) results are in fact applicable to our framework with firms differentiated in quality instead of productivity.

[36]The social planner chooses to equalize markups across firms at $m = \frac{\kappa-\gamma}{\kappa-\gamma-1}$.

[37]We note, the quality standard is not first-best. It raises expected profits, which induces too much entry, and cannot bring the economy to the efficient allocation.

is absent, the average markup in the economy increases because of the composition effect.

DA Preferences. Under DA preferences ($\eta \to \infty$), the market allocation generates business stealing bias, provided that $\gamma > 0$ and, thus, demand is not fully rigid. The ratio of the planner's cutoff to the market cutoff is $\left(\frac{z_P^*}{z^*(1)}\right)_{DA} = \left(1 + \frac{1}{\gamma}\right)^{\frac{\gamma}{\kappa-\gamma}} \geq 1$. For $\gamma > 0$, a quality standard improves welfare, by reducing the business stealing bias. For $\gamma = 0$, the bias disappears and the standard cannot improve welfare. The main difference relative to the IA case is that a standard has anti-competitive effect under DA preferences: as the standard reduces the number of firms in the market, the lower competitive pressure allows for surviving firms to charge higher markups, limiting the benefits of the standard.

Homothetic Preferences. Under homothetic preferences ($\eta = 0$), the ratio of the planner's cutoff relative to the market cutoff is $\left(\frac{z_P^*}{z^*(1)}\right)_H = \left(1 + \frac{1}{\gamma}\right)^{\frac{\gamma}{\kappa-\gamma-1}} \left(1 - \frac{1}{\kappa-\gamma}\right)^{\frac{1}{\kappa-\gamma-1}}$, and it could be smaller or greater than one depending on the parameters of the model[38]. If the ratio is greater than one, the effects of a standard are qualitatively similar to the DA case.

In the presence of a dominating appropriation bias, a quality standard can still improve welfare, although by a somewhat smaller magnitude relative to the case in which there is too little selection. The reason for this seemingly surprising result is that the market allocation generates a markup distribution that is different from the constant markup that a planner would choose. In particular, markups are on average too small in the market relative to the planner's allocation. The quality standard improves upon such misallocation, despite exacerbating the already too high level of selection. Therefore, one important conclusion from our analysis is that the market distortions are driven entirely by the presence of variable markups, and exist in both homothetic and non-homothetic preferences.

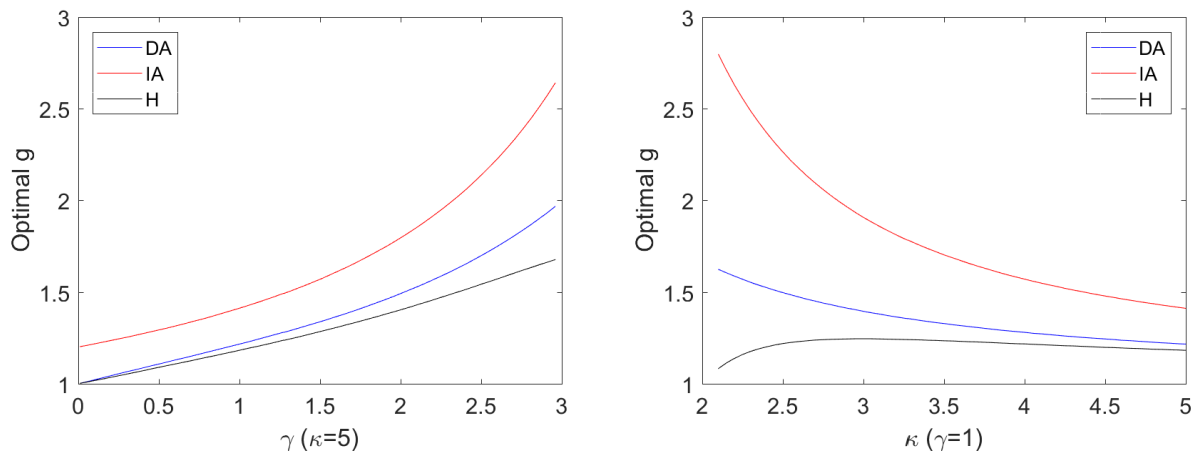
Optimal Standard Across Preferences. Figure 2 shows a ranking of the optimal degree of restrictiveness of the standard g across preferences. In particular, $g_{IA}^{opt} > g_{DA}^{opt} > g_H^{opt}$. The reason for this ranking is related to the extent of the anti-competitive effects of the standard, which depends on the elasticity of the market cutoff with respect to market size $\frac{\partial \ln z^*}{\partial \ln L}$. Under IA preferences, such an elasticity is zero and the anti-competitive effect are absent. Thus, the optimal standard is the most restrictive. Under DA and homothetic preferences, the reduction in the number of firms generate increases in markups. Such increases are the

[38]The business stealing bias dominates, if $\kappa > \gamma + \left(1 - \left(1 + \frac{1}{\gamma}\right)^{-\gamma}\right)^{-1}$. Since regularity conditions imply that $\kappa > \gamma + 1$, there is a region for small enough κ, in which appropriation bias dominates. For instance, for the linear case $\gamma = 1$, there is too much selection if $\kappa \in (2, 3)$. Such a case is not quantitatively relevant: in the empirical section we verified that it only occurs in one industry.

largest under homothetic preferences. Hence, the optimal standard is the smallest under homothetic preferences, and it is at an intermediate level under DA preferences[39].

Figure 2: Optimal g Across Preferences

(a) As a function of γ **(b)** As a function of κ



3.5 The Effect of Trade on the Optimal Standard

The purpose of this section is to study whether international trade modifies the optimal level of a quality standard. We consider a world made of two identical economies, in which exports require an iceberg trade cost τ. As the two countries are identical, we can normalize per capita income and wages to one in both countries. We assume that the two countries decide a common minimum quality requirement $\bar{z}$ and we study how different levels of the iceberg trade cost τ affect the optimal level of $\bar{z}$. Thus, we abstract from political economy issues — as well as a complicated non-cooperative game — that would arise if each country were to choose its own optimal level of $\bar{z}$.

As in a standard Melitz (2003) model, the presence of iceberg trade costs divides firms into exporters and non-exporters. Only high-quality firms are exporters, since their quality level is large enough to sell their goods abroad. Details are in Appendix 6.2.6. Let z_D^* and $z_X^* = \tau z_D^*$ denote the market quality cutoff for the domestic and export market. Similarly to the closed economy case, let $g = \frac{\bar{z}}{z_D^*}$ be a measure of the restrictiveness of the quality standard. The effect of g on the selection of firms exhibits a discontinuity. In particular, for $g \in [1, \tau]$, the quality standard only affects low-quality non-exporters. On the other hand,

[39]The ranking of optimal g as a function of the degree by which markups depend on the number of competitors is respected across other preferences not included in GTP. In the online appendix, we provide a detailed discussion of the (IA) addilog preferences (Bertoletti et al., 2018), (DA) Stone-Geary (Simonovska, 2015), and (homothetic) Quadratic Mean of Order R (Feenstra, 2018). Finally, we explore the effects of variety externality in the Benassy-CES preferences (Benassy, 1996)

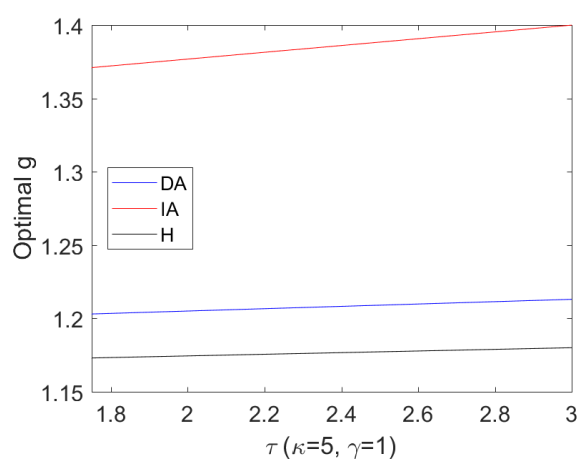
for $g \in [\tau, \infty)$, the standard is so restrictive that the only firms able to remain active are high-quality exporters.

How does the level of international openness affect the optimal minimum quality requirement? To answer to this question, we compute the optimal g^{opt} for different levels of τ. As shown in Figure 3, lower trade costs reduce the optimal quality requirement[40].

Under homothetic and DA preferences, a reduction in trade costs increases the market selection in the domestic economy without a standard, namely z_D^* increases. Trade forces some of the firms of low-quality out of the market, in a similar fashion to the quality standard. Hence, a reduction of trade costs and an increase in the restrictiveness of the standards are two policies that achieve the same welfare result. Therefore, the larger the reduction in trade costs, the larger the reduction in the restrictiveness of standards.

Under IA preferences, the domestic cutoff z_D^* is not affected by trade costs: a trade liberalization does not generate selection effect that drives domestic firms out of the market. However, it is reassuring that even in this extreme case, lower iceberg trade costs still improve allocative efficiency. The rationalization of this finding comes through production reallocation. As trade costs decline, export production increases. Although there is the same number of domestic varieties available, these now compete with foreign varieties, which reallocates production towards high-quality goods. Therefore, the level of distortions is *lower* in a more open economy, alleviating the need for a standard. We believe this rationalizes a dual approach for policymakers: pushing towards lower trade costs while eliminating unnecessary restrictiveness of quality standards.

Figure 3: Optimal Requirement and Trade Costs



[40]We restrict τ so that $g^{opt} \in [1, \tau]$, and the restriction only affects non-exporters. Although it is theoretically possible for the standard to force out of the market all non-exporters, it is not a realistic outcome and, thus, we ignore it.

4 Model Estimation

In the theoretical framework, we incorporate technical measures into an economy with firm heterogeneity, so that the measure of producing firms and the sales of these firms depend on the level of regulatory restrictiveness in the industry plus demand and supply parameters. Next, we use the model to estimate the regulatory restrictiveness of Chilean industries by matching the empirical sales distribution constructed with the firm data described in Section 2.

We do not require any data on the number of standards imposed, but instead employ a simulated method of moments (SMM) procedure that estimates (g, κ, γ) to minimize the difference between percentiles of the model and data sales distribution. We run this procedure for a cross-section of industries and repeat it across multiple years. The estimation yields not only an implied level of restrictiveness – which we call g in the model – but also the optimal level of restrictiveness at the industry level given the supply and demand parameters. Hence, we provide a meaningful interpretation of how many industries appear to be *too restrictive* as characterized by the structure of the model, even allowing for large optimal standards. Previewing the results, we find that many industries appear very restrictive up until 2000, but not in 2005.

4.1 Strategy

In this section, we describe how to quantify the restrictiveness in a market by using the structural model described in Section 3. Our goal is to estimate the parameter set (g, κ, γ) for each industry, as these are enough to characterize the "restrictiveness" of an industry as given by g/g^{opt}. We solve the model via simulation because the moments in the model that pin down these parameters are created using simulated firms. In other words, for a guess of the parameters, we simulate firm-level outcomes and attempt to reproduce moments from the empirical domestic sales distribution.

First, we simulate a large enough number of draws so as to best approximate the entire continuum of firms that exist in the model. We follow the insights of Eaton et al. (2011) that were recently applied in Jung et al. (2019), and relabel firm-level indicators that can be simulated from a parameter-free uniform distribution. Recall that the pdf of the quality distribution is given by $h(z) = \frac{\kappa b^{\kappa}}{z^{\kappa+1}}$. We draw 500,000 realizations of the uniform distribution on the [0; 1] domain, $U \sim [0; 1]$, we order them in increasing order, and find the maximum realization, denoted by u_{max}. Then, the firm quality indicator is $z = (u/u_{max})^{-1/\kappa} z^*$. Given that the market quality cutoff (z^*) is a constant in the IA case, we normalize this to one[41].

[41] In the non-IA cases, we compute the new z^* for any given guess of the standard.

By construction, $z \in [1, \infty]$, and hence all draws have positive demand in the case where there is no government imposed quality standard. Given that there exists restrictions on the survival of low-quality firms, the set of producing firms is chosen from $z \in [g, \infty]$.

We adopt an over-identification strategy that targets 99 moments from the empirical domestic sales distribution. Given a set of potential producers in the simulation, namely those with $z > g$, we compute firm revenues normalized by mean revenues:

$$\tilde{r}(z|z > g) = \frac{r}{\bar{r}} = (G_2(g))^{-1} \left(\frac{z}{z^*} - 1\right)^{\gamma} \left(\frac{z}{z^*} + \gamma\right) \tag{17}$$

where $F_1(g)$ and $F_2(g)$ are two hypergeometric functions described in footnote 29. After conditioning on active firms, relative sales are independent of η[42].

The theoretical relative sales are matched to their counterpart in the data in order to identify the model parameters in an approach that follows Sager and Timoshenko (2017). Let $F_q^m(g, \kappa, \gamma) = \log(\tilde{r})_q$ be the q-th quantile of the simulated log domestic sales distribution. Then, let F_q^d denote the corresponding value of the empirical CDF of the log sales distribution. Our identification consists of choosing the parameter set that minimizes the sum of the squared errors between empirical and theoretical quantiles:

$$\min_{g,\kappa,\gamma} \sum_{q=1}^{99} \left(F_q^d - F_q^m(g, \kappa, \gamma)\right)^2 . \tag{18}$$

Finally, we compute bootstrap standard errors by running the estimation above 100 times, each time taking a bootstrap sample of the data. We take the average parameter estimates $(\hat{g}, \hat{\kappa}, \hat{\gamma})$, and use the standard deviation of estimates to compute a 95% confidence interval. We employ the strategy above for each 4-digit ISIC (revision 3) industry by year. Although there are about 100 industries in this level of classification, we only keep those with at least 35 firms in 1995, which allows us to estimate restrictiveness for 40 industries.

The strategy to estimate the parameter set $(\hat{g}, \hat{\kappa}, \hat{\gamma})$ is based on the separate ways that each parameter is identified within the sales distribution. κ governs the shape of the quality distribution, which is proportional to the shape in the sales distribution only in special cases (Mrazova et al., 2017), which do not apply to our GTP specification. The divergence in the sales and quality distribution is due to the distribution of markups. Since firm markup levels are a function of γ (see (8)), this parameter affects the mapping from the quality to the sales distribution and is not collinear with κ[43]. Finally, as is argued above, the standard not

[42] In general, z^* depends on η (11). However, the estimated parameters across the types of preferences are almost identical. Of course, the predicted optimal standard depends on η, and in the appendix we compare the results between the IA and DA cases.

[43] As is not the case, for example, if preferences were CES and the distribution of quality is Pareto.

only eliminates low-quality firms but reallocates resources to higher-quality firms. Therefore, relative sales across percentiles of the sales distribution are a function of g. For this reason we use a general strategy to match sales across the firm distribution, with each parameter being identified by different parts of the distribution.

4.2 Estimation Results

For expositional purposes, we employ the procedure outlined in the previous section to estimate $(\hat{g}, \hat{\kappa}, \hat{\gamma})$ for the universe of Chilean manufacturing firms in each year. The estimated level of the quality standard $\hat{g}$ had a 95% confidence interval above one in every year, with the standard peaking at 1.13 (with standard error of .01) through 1998-2000, before dropping every year thereafter to 1.02 in 2007 (standard error .002).

Panel A of Table 2 displays the results for the parameter estimates in 1995, 2000, and 2005. In Panel B, we report the data value and the simulated value for 5 moments that are indirectly targeted[44]. The parameter for demand curvature $\hat{\gamma}$ ranges between 1.3 and 2.4, rejecting the simple linear demand model in every year. The Pareto shape parameter of the quality distribution $\hat{\kappa}$ varies between 4-5 for the majority of the sample (consistent with estimates in Jung et al. (2019) and Simonovska and Waugh (2014)), and below 3 after 2004. It is evident that we can match moments from the sales distribution closely. Finally, the model implied average markups are 14%, 12%, and 31% in the three displayed years. These are very much in line with the empirical average markup estimates, which are not targeted in the estimation[45].

The empirical sales distribution exhibits structural changes captured by the variation of $\hat{\kappa}$ and $\hat{\gamma}$ over time. Although $\hat{\kappa}$ is roughly constant between 1995 and 2003, the subsequent reduction in $\hat{\kappa}$ highlights an increase in the underlying quality dispersion. There are several moments of the sales distribution that help explain the observed movement in the parameters of the model. For example, we find that the ratio of sales between exporters and non-exporters gets larger[46], differences in sales between firms at certain percentiles become larger, skewness decreases (longer left tail), and average markups increase (as reported above). This could reflect an expansion of the largest firms, or smaller firms at the bottom of the distribution. We find that the former is more apparent in the sales data, and point to the effect of trade in the industry analysis below.

[44]In the appendix, Figure 12 displays the model and empirical sales distributions, which allows us to visually compare the model and empirical sales distributions, which are reassuringly close.

[45]These are based on the De Loecker and Warzynski (2012) procedure and using material inputs as the variable input. We take a weighted average using the firms' share of total employment in the economy.

[46]This holds with average sales of firms above the median in domestic sales versus below the median, which are reported in Table 2.

Table 2: Estimation Results: Manufacturing-wide in 1995, 2000, and 2005

Panel A: Parameter Estimates

Year	Data Targets	$\hat{g}$	$\hat{\kappa}$	$\hat{\gamma}$
1995	Sales Percentiles (1-99)	1.1 (.01)	4.49 (.57)	1.92 (.15)
2000	Sales Percentiles (1-99)	1.13 (.01)	4.94 (.69)	2.40 (.18)
2005	Sales Percentiles (1-99)	1.02 (.004)	2.45 (.17)	1.30 (.06)

Panel B: Moments: Data versus Model

	1995		2000		2005	
Moment	Data	Model	Data	Model	Data	Model
Sales Advantage	2.41	2.42	2.57	2.59	2.88	2.87
90-10 Sales	3.89	3.88	4.15	4.15	4.71	4.65
99-90 Sales	2.02	2.04	2.25	2.28	2.26	2.35
Skewness	.67	.69	0.78	.79	.43	.33
Average Markup	22%	14%	20%	12%	38%	31%

Panel A reports the parameter estimates when estimating the model for all manufacturing firms in each year (results also available per industry, but are not reported as they would contain 40 estimates for each parameter per year). We compute bootstrap standard errors (in parenthesis) by running the estimation 100 times, each time taking a bootstrap sample of the data. In Panel B, we compute 5 moments in the data and using the simulated firms. "Sales Advantage" reflects the log difference between the average sales of firms in the to 50% relative to the bottom 50%. "90-10" and "99-90" are log differences in sales between firms in the respective percentiles. The average markup in the data is computed from estimating markups using the De Loecker and Warzynski (2012) procedure and taking a weighted average using the firms' share of total employment in the economy.

The estimated restrictiveness of the standard $\hat{g}$ is highly heterogeneous across industries. Consider the reduction in the survival probability due to the standard, which is calculated as $(1 - \hat{g}^{-\hat{\kappa}})$. Table 3 reports the reduced probability for each industry in 1995, 2000 and 2005. The most restrictive industries across the years – Motor Vehicles, Books/Journals, Machinery and Other Metals, Pharmaceutical, and various food industries – averaged over 50% lower survival probability. On the other hand, there are industries such as Apparel and Furniture that hover around 10% in reduced probability across all years[47]. On average, the reduction in probability is 37% in 1995, 43% in 2000, and drops to 29% in 2005. This is consistent with the previous results, where the restrictiveness at the economy wide-level seems to decrease in 2005.

Robustness. The estimation above takes a general approach in terms of attempting to match the whole sales distribution instead of specific moments within the distribution. As a robustness check, we have applied a similar SMM procedure with specific moments from the sales distribution that are pinned down by our parameters of interest. We construct 4

[47]In the analysis of time variation below, we eliminate "Other Publishing" and "Bakery" which have massive decreases in restrictiveness. There are also a few industry-year pairs we drop from the table as the estimated $\hat{\kappa}$ is imprecise.

moments: i) the sales advantage of "high-quality" relative to "low-quality" firms[48]; ii) the skewness of the distribution which captures the composition effect of the standard; and two differences: iii) $\log(\tilde{r})_{99} - \log(\tilde{r})_{90}$, and iv) $\log(\tilde{r})_{90} - \log(\tilde{r})_{10}$[49]. We plot the simulated sales distribution in Appendix 6.3 and report the estimated parameters. We do not find large discrepancies with our benchmark strategy: $\hat{g}$ is very similar and has the same time series patterns, as do the other two parameters. However, in the alternative calibration, the fit with the data distribution is clearly not as close.

As an alternative approach, we consider fixing some of parameters to estimate restrictiveness, with full results in Appendix 6.3. First, we set $\gamma = 1$ and estimate the remaining two parameters, in case there is collinearity with the shape of the quality distribution. A graphical analysis shows that this specification is not able to match the dispersion in sales in the data. Moreover, the implied average markups, which are 39%, 46%, and 51%, are larger than reasonable markup estimates. However, the time series pattern of $\hat{g}$ uncovered in the baseline estimation holds[50]. Second, we estimate *only* the restrictiveness parameter, and set $\kappa = 4$ and $\gamma = 1.8$ as deep parameters constant over time. Still, $\hat{g}$ ranges between 1.04 and 1.1 with similar time-series movements[51].

4.3 Restrictiveness of Standards and Optimal Standard

We compare the estimated level of restrictiveness with the optimal standard predicted by our model. We construct a restrictiveness index (RI) using the estimated parameters:

$$RI_{it} = \frac{\hat{g}_{it}}{\hat{g}_{it}^{opt}(\hat{\kappa}_{it}, \hat{\gamma}_{it})}. \tag{19}$$

where i denotes an industry, and t a year. The interpretation of this index is different from the technical measures in Section 2, as it captures a wide variety of measures – for example one that is meant to be protectionist in the guise of a quality standard – that limits the

[48]This moment is related to that used to identify the elasticity of substitution in Bernard et al. (2003). Instead of comparing exporters and non-exporters, we compare firms above and below the median in sales.

[49]We do not use moments from other distributions, such as markups and value added, because the data does not allow us to differentiate between exported and domestic components.

[50]However, there is a shift downwards of $\hat{\kappa}$ and $\hat{g}$, as the restriction on the demand curvature requires a larger dispersion in quality to match the dispersion in sales. The result supports the findings of Jung et al. (2019), who estimate the Melitz and Ottaviano (2008) separable model and argue that the linear demand assumption predicts a sales distribution with too little dispersion (given their preferred estimate of κ as estimated using a general demand curvature). Relative to the linear case, a larger γ raises the sales of the largest firms as demand is more elastic.

[51]As a further robustness check, we have estimated the addilog model of Bertoletti and Etro (2017) and the linear, separable Melitz and Ottaviano (2008) model. The former is very similar to the IA case in our GTP framework, while the latter is nested in our DA case with $\gamma = 1$. We found the results are almost identical to those two estimations, and we make the results available upon request.

survival of firms at the bottom of the sales distribution. We choose the IA model in a closed economy as the benchmark because it yields the most conservative estimate of whether an industry is *too* restrictive[52]. We interpret g_{IA}^{opt} as an *upper bound* for which policymakers can view an industry as overly regulated. For each industry-year, we compute (19) and find that there are several industries that appear too restrictive but that number has changed over the years. As in the estimation for the universe of firms, the level of restrictiveness increases from 1995 to 2000, but drops significantly in 2005.

Figure 4 plots the RI in 1995, 2000, and 2005 for each industry, sorted from largest to smallest. For each industry-year, we derive the 95% confidence interval using the estimated standard error for g in the calibration. We define industries with a confidence interval for RI that includes a ratio of one or above as too restrictive. In 1995, 11 out of 38 industries are too restrictive, although there are several industries that hover around 1[53]. In 2000, there are 12 too-restrictive industries, though a similar number are clustered around an RI index of 1. Therefore, even with the conservative measure of the optimal standard, 32% of industries are within the confidence interval of being too restrictive. We take this as evidence that Chilean manufacturing industries appear to be overly regulated – either through protectionism or other types of regulations – at the start of the century. However, this restrictiveness drops precipitously over the next few years. In 2005 only 5 out of 38 industries (13%) were overly regulated, and many more industries drop far below the cutoff[54].

Although this estimation does not allow us to break down the restrictiveness into specific measures, it is likely that a greater openness to trade contributed to the reduction. In the appendix (Figure 19), we plot the total value of exports and imports relative to GDP, which suggests that the economy becomes more open after 2000. Furthermore, Chile signed free trade agreements with the EU in 2002, with the United States and Korea in 2004, and with China in 2005. It also lowered its across-the-board tariffs to 6% for all countries with which it did not have an agreement. In the bottom panel of Figure 19, we plot average applied tariff rates (weighted by industry import flows) and the terms of trade (provided by the World Development Indicators). Tariffs begin their decline in 1999, dropping from 11% to 2% in 2007. There is a large terms of trade appreciation after 2001 – due to the price of copper – which would create opportunities for importers to enter the Chilean market. We find that there is a clear negative correlation between the changes in restrictiveness and openness of

[52]Trade reduces g^{opt} (see section 3.5). Under DA and homothetic preferences, more industries would appear too restrictive as the implied g^{opt} is smaller (results in the appendix).

[53]The noise in the estimation can affect whether an industry fits within our definition of too restrictive. However, this is only obvious in the "Other Manufacturing" and "Journals" industries.

[54]Plots that result from the estimation of DA preferences and assuming a fixed cost for the standard are in the appendix – in both cases the optimal standard is merely shifted downwards.

the industry[55].

Recall that freer trade provides a welfare-enhancing reallocation to high-quality firms, diminishing the necessity of the standards. For example, take the meat industry (ISIC 1511), one of the most open industries in Chile. The ratio of average domestic sales of exporters relative to average domestic sales of non-exporters is 2.80 in 2000, and increases to 3.44 in 2005[56]. The results suggest that the expansion of the large firms is what drives the estimated lower restrictiveness of industries (and is also consistent with a lower estimated κ). Although the contemporaneous correlation is only suggestive evidence, it does reinforce our claim that the rationale introduced in this paper supports a dual approach of lowering trade costs while reducing unnecessary restrictiveness of quality standards.

5 Conclusion

We have provided a rationale for regulations on product standards that does not rely on the existence of negative externalities nor the benefits of quality upgrading. The main contribution of this paper is to provide a framework under which standards can improve welfare by reallocating production from low-quality firms to high-quality firms. In order to motivate the welfare-enhancing reallocation that occurs in the model, we rely on a panel of Chilean manufacturing firms and compare the distribution of firm sales across industries that differ in their level of regulation. Our findings that technical measures skews domestic sales towards high-quality firms complements the rest of the literature that has found these measures to reduce the extensive margin of export flows.

Moreover, we have shown theoretically that trade openness reduces the optimal degree of restrictiveness of quality standards. The theoretical prediction finds support in the data: openness appears to have reduced the regulatory burden in Chilean manufacturing. We estimate the model to fit the observed distribution of domestic sales and conduct a policy-relevant evaluation that compares the estimated level of restrictiveness with the optimal standard as predicted by our model. Although industries appear heavily regulated up until 2000, this is not the case in 2005 and there is suggestive evidence that it is driven by more open industries. Hence, this paper supports the efforts of the WTO to improve the Technical Barriers to Trade Agreement, along with traditional trade agreements.

[55] Figure 20 in the appendix plots the log difference in RI between 2000 and 2005 against an openness measure defined as the sum of imports and exports over total sales. The correlation is -.20.

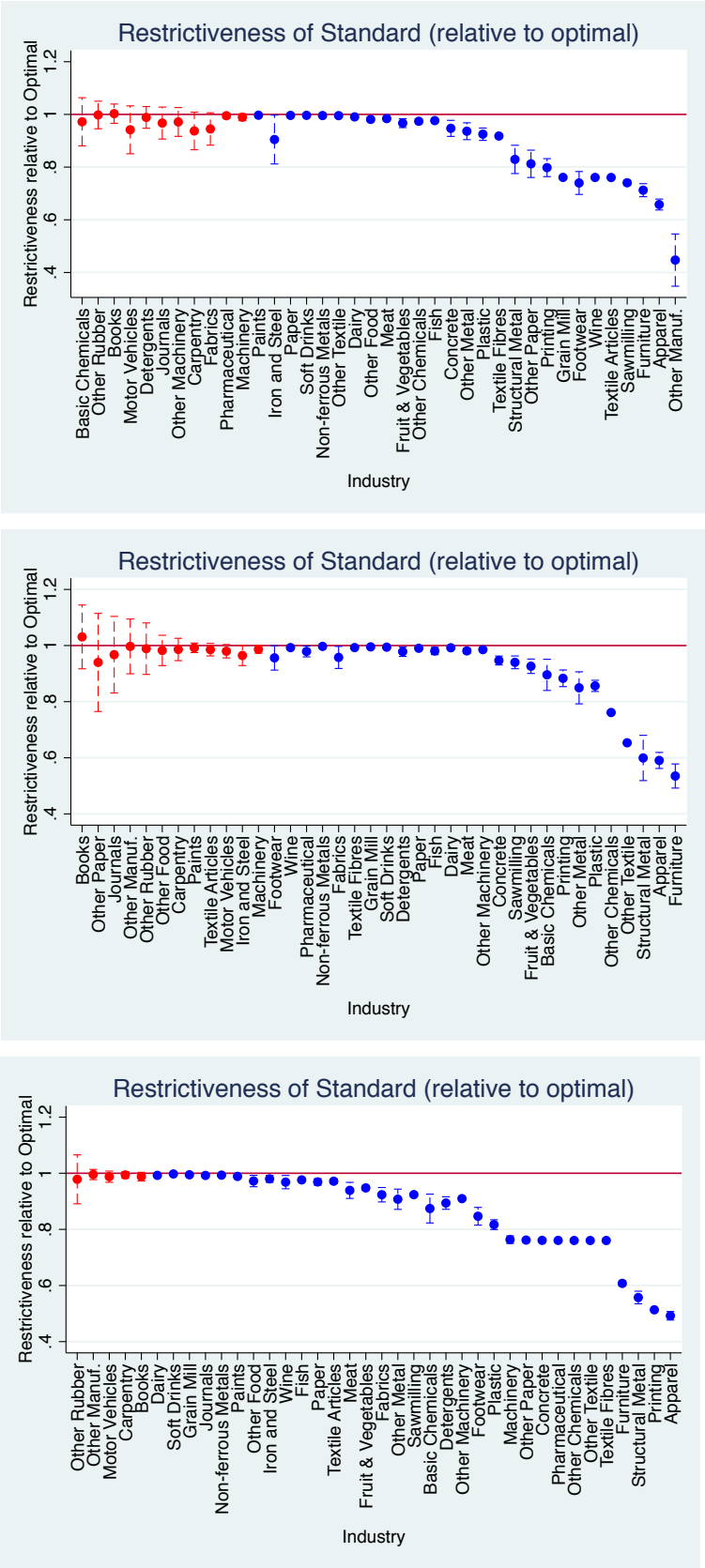
[56] The values are similar for all years between 1995-2003, and 2004-2007. Furthermore, a similar pattern holds across many open industries, related to the "Sales Advantage" in Panel B of Table 2.

Table 3: Reduction in Probability of Survival Due to Restrictions by Industry-Year

ISIC	Industry name	1995	2000	2005
1511	Meat	0.4	0.45	0.47
1512	Fish	0.16	0.38	0.2
1513	Fruit & Vegetables	0.35	0.33	0.13
1520	Dairy	0.32	0.22	0.3
1531	Grain Mill	-	0.06	0.08
1549	Other Food	0.55	0.72	0.41
1552	Wine	-	0.3	0.42
1554	Soft Drinks	0.12	0.33	0.1
1711	Textile Fibres	0.04	0.25	-
1721	Textile Articles	-	0.49	0.13
1729	Other Textile	0	0	0.8
1730	Fabrics	0.59	0.57	0.25
1810	Apparel	0.12	0.15	0.02
1920	Footwear	0.32	0.61	0.17
2010	Sawmilling	0.06	0.38	0.14
2022	Carpentry	0.31	0.71	0.48
2102	Paper	0.23	0.35	0.31
2109	Other Paper	0.22	0.84	0
2211	Books	0.87	0.98	0.45
2212	Journals	0.71	0.81	0.45
2221	Printing	0.2	0.27	0
2411	Basic Chemicals	0.8	0.43	0.21
2422	Paints	0.09	0.49	0.22
2423	Pharmaceutical	0.48	0.42	-
2424	Detergents	0.75	0.65	0.11
2429	Other Chemicals	0.19	-	-
2519	Other Rubber	0.79	0.78	0.58
2520	Plastic	0.37	0.26	0.19
2695	Concrete	0.5	0.26	-
2710	Iron and Steel	0.47	0.72	0.39
2720	Non-ferrous Metals	0.2	0.13	0.11
2811	Structural Metal	0.46	0.22	0.04
2899	Other Metal	0.49	0.48	0.48
2919	Other Machinery	0.62	0	0
2924	Machinery	0.43	0.53	-
3430	Motor Vehicles	0.6	0.61	0.56
3610	Furniture	0.15	0.13	0.03
3699	Other Manuf.	0.16	0.84	0.6
	Average	0.37	0.43	0.29

This table reports the reduced probability of producing in each industry given the estimated restrictiveness of the industry. The probability is calculated as $(1 - \hat{g}^{-\hat{R}})$. It is based on the simulated method of moments estimation for year industry by year. For certain industry-year pairs, the data was insufficient to estimate stable parameter values, which is why certain entries above are missing.

Figure 4: Restrictiveness Index with IA Preferences: 1995, 2000, and 2005.



In red are the industries where we cannot reject the null hypothesis that the restrictiveness index is different from one.

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6 Appendix

6.1 Motivational Evidence

Table 4: Top 25 Most Regulated Industries

	Ranked by Coverage		Ranked by Total Number	
Rank	ISIC	Industry Name	ISIC	Industry Name
1	2421	Pesticides	2411	Basic chemicals
2	1520	Dairy products	1512	Fish products
3	1531	Grain products	1511	Meat products
4	1552	Wine	1711	Textiles
5	1511	Meat products	1513	Fruit and vegetables
6	1513	Fruit and vegetables	1810	Wearing apparel
7	1551	Alcohol production	2423	Pharmaceuticals
8	1554	Soft drinks	1514	Oils and fats
9	1532	Starch products	1520	Dairy products
10	1533	Animal feeds	1549	Other Food
11	1549	Other Food	1531	Grain products
12	1512	Fish products	2922	Machine tools
13	1514	Oils and fats	2429	Other chemicals
14	2424	Cleaning products	2930	Domestic appliances
15	2010	Wood	2520	Plastic
16	1544	Farinaceous products	3210	Electronic components
17	1543	Candy bars	2424	Cleaning products
18	2021	Plywood, etc	2919	Other general purpose machinery
19	3230	TV and radio receivers	3190	Other electrical equipment
20	3150	Lighting equipment	3230	TV and radio receivers
21	3190	Other electrical equipment	3110	Electric motors
22	2912	Pumps	1532	Starch products
23	3311	Medical equipment	2010	Wood
24	2423	Pharmaceuticals	3311	Medical equipment
25	2023	Wooden containers	2912	Pumps

Ranked by Coverage refers to the standard approach of normalizing the number of regulations in an industry by the number of products in the industry. We compare this with the total number of regulated products in an industry (Ranked by Total Number).

Table 5: Correlation of Quality with TFP (top) and Size (bottom) across Firms

	Quality Proxy					
	log(K/L)	log(W/L)	log(M/L)	log(K/L)	log(W/L)	log(M/L)
Log TFP	0.532***	0.436***	1.323***			
	(0.110)	(0.036)	(0.068)			
Log Labor Productivity				0.646***	0.349***	0.606***
				(0.030)	(0.012)	(0.030)
Fixed Effects	Industry-Year	Industry-Year	Industry-Year	Industry-Year	Industry-Year	Industry-Year
# Observations	63790	63785	63790	61779	65483	65441

	Quality Proxy		
	(K/L)	(W/L)	(M/L)
Log Size	0.464***	0.227***	0.473***
	(0.015)	(0.005)	(0.014)
Fixed Effects	Industry-Year	Industry-Year	Industry-Year
# Observations	64894	68864	67993

This table regresses the quality proxy on measures of productivity and size. The top table displays results for two different measures of productivity: log TFP estimated from a Translog production function using the procedure outlined in Weinberger (2017), and a simple measure of logged value added per worker. The bottom panel displays results for the relationship between quality and log sales. Quality is proxied by (logged) capital per worker, total wages per worker, and input expenditure per worker respectively. In all specifications we use the full panel of firm-year observations and include industry-year fixed effects so that we are only capturing within industry-year relationships. Standard errors – clustered by 4-digit industry – are in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 6: Firm Sales and Survival Heterogeneity - by TMs in Industry: All Controls Shown

	Log Domestic Sales			Survival		
	(1)	(2)	(3)	(4)	(5)	(6)
TM*Quality (Capital/L)	0.015**			0.012**		
	(0.006)			(0.005)		
Openness*Quality (Capital/L)	0.965			-0.526***		
	(0.787)			(0.099)		
Tariff*Quality (Capital/L)	-0.011***			-0.008***		
	(0.004)			(0.002)		
TM*Quality (Wage/L)		0.012**			0.009	
		(0.006)			(0.006)	
Openness*Quality (Wage/L)		0.144			-0.406	
		(1.094)			(0.271)	
Tariff*Quality (Wage/L)		-0.012***			0.000	
		(0.004)			(0.001)	
TM*Quality (InputValue/L)			0.008			0.006
			(0.006)			(0.004)
Openness*Quality (InputValue/L)			-1.998**			-0.206
			(0.769)			(0.129)
Tariff*Quality (InputValue/L)			-0.006			-0.001
			(0.005)			(0.001)
Fixed Effects	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y
R^2	0.955	0.955	0.955	0.645	0.644	0.643
# Observations	44220	44220	43789	69679	69679	68924

Results from the same specification as Table 1, but here we display results for all control variables. TM_{it} (restrictiveness) is measured at the 4 digit ISIC industry level. The total number of measures in each industry-year are summed and then divided by the number of HS6 products in the industry. Each row interacts the TM measure with a dummy for quality, where quality is proxied by capital per worker, total wages per worker, and input expenditure per worker respectively. Firms are high-quality if the quality proxy is above the median in that industry in 1995. For the results on survival, all firms alive in 1995 are "potential" producers in all years, which is why the number of observations is much larger. In all specifications we include an interaction of industry openness with the quality indicator, an interaction of the quality indicator measure with the industry import tariff, plus firm and industry-year interacted fixed effects. Standard errors – clustered by 4-digit industry – are in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 7: Firm Sales Heterogeneity - by NTMs in Industry - Repeated Cross-Sections

	Domestic Sales		
	(1)	(2)	(3)
TM*Quality (Capital/L)	0.037		
	(0.039)		
Openness*Quality (Capital/L)	-2.856		
	(2.164)		
Tariff*Quality (Capital/L)	-0.049***		
	(0.012)		
DemandElast*Quality (Capital/L)	-0.025		
	(0.026)		
TM*Quality (Wage/L)		0.006	
		(0.040)	
Openness*Quality (Wage/L)		-3.538	
		(2.426)	
Tariff*Quality (Wage/L)		-0.037***	
		(0.008)	
DemandElast*Quality (Wage/L)		-0.028	
		(0.027)	
TM*Quality (InputValue/L)			0.049
			(0.043)
Openness*Quality (InputValue/L)			-4.136*
			(2.390)
Tariff*Quality (InputValue/L)			-0.050***
			(0.008)
DemandElast*Quality (InputValue/L)			-0.032
			(0.027)
Fixed Effects	Industry-Year	Industry-Year	Industry-Year
R^2	0.362	0.446	0.465
# Observations	59740	59740	59740

This Table is a repeated cross-section analysis of specification (1). Regulations are aggregated by industry across all years, so variation is only at the industry level. TM_i (restrictiveness) is measured at the 4 digit ISIC industry level. The total number of measures in each industry are totaled and then divided by the number of HS6 products in the industry. Each row interacts the TM measure with a different proxy for quality. In all specifications we include an interaction of industry openness with the quality indicator, an interaction of the quality indicator measure with the industry import tariff, and an interaction of the quality indicator measure with the industry trade elasticity (Broda and Weinstein (2006)). We include only industry-year interacted fixed effects. Standard errors – clustered by 4-digit industry – are in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 8: Firm Sales and Survival Heterogeneity- by NTMs in Industry - IV Results

	Domestic Sales			Survival		
	(IV)	(IV)	(IV)	(IV)	(IV)	(IV)
TM*Quality (Capital/L)	0.034**			0.002		
	(0.016)			(0.011)		
TM*Quality (Wage/L)		0.046***			0.031***	
		(0.016)			(0.011)	
TM*Quality (InputValue/L)			0.006			0.002
			(0.016)			(0.011)
F-stat (first stage)	1745.11	1731.37	1744.89	1862.60	1862.60	1864.97
# Observations	44220	44220	43789	69679	69679	68924
Fixed Effects	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y

In this Table we run a 2SLS IV variant of (1). In each case, the main interaction is instrumented using the interaction of the same quality indicator but with the TM measure of Peru. The measure of TMs is at the 4 digit ISIC industry level. The total number of measures in each industry-year are totaled and then divided by the number of HS6 products in the industry. Each row interacts the TM measure with a different proxy for quality. In all specifications we include an interaction of industry openness with the quality indicator, an interaction of the quality indicator measure with the industry import tariff, plus firm and industry-year interacted fixed effects. Standard errors – clustered by 4-digit industry – are in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 9: Firm Sales and Survival Heterogeneity - by specific types of TMs: SPS (top), TBT (middle), Non-Domestic (Bottom)

	Log Domestic Sales			Survival		
	(K/L)	(W/L)	(M/L)	(K/L)	(W/L)	(M/L)
TM*Quality (Capital/L)	0.015***			0.008*		
	(0.005)			(0.004)		
TM*Quality (Wage/L)		0.009*			0.011	
		(0.006)			(0.007)	
TM*Quality (InputValue/L)			0.006			0.007
			(0.006)			(0.005)
Fixed Effects	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y
R^2	0.955	0.955	0.955	0.645	0.644	0.643
# Observations	44220	44220	43789	69679	69679	68924

	Log Domestic Sales			Survival		
	(K/L)	(W/L)	(M/L)	(K/L)	(W/L)	(M/L)
TM*Quality (Capital/L)	0.010			0.026**		
	(0.019)			(0.011)		
TM*Quality (Wage/L)		0.023			0.012	
		(0.019)			(0.012)	
TM*Quality (InputValue/L)			0.016			0.003
			(0.022)			(0.009)
Fixed Effects	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y
R^2	0.955	0.955	0.955	0.645	0.644	0.643
# Observations	44220	44220	43789	69679	69679	68924

	Log Domestic Sales			Survival		
	(K/L)	(W/L)	(M/L)	(K/L)	(W/L)	(M/L)
TM*Quality (Capital/L)	0.042			-0.004		
	(0.026)			(0.012)		
TM*Quality (Wage/L)		-0.007			-0.000	
		(0.030)			(0.015)	
TM*Quality (InputValue/L)			-0.009			-0.011
			(0.025)			(0.013)
Fixed Effects	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y	Firm, I-Y
R^2	0.955	0.955	0.955	0.645	0.644	0.643
# Observations	44220	44220	43789	69679	69679	68924

In this table we conduct the specification displayed in (1), but *for specific types of Non-Tariff Measures*. In the top panel we allow technical measure for the SPS chapter only, but drop those geared towards imports. In the middle panel we allow technical measure for the TBT chapter only, but drop those geared towards imports. In the bottom panel we include "A1" and "B1" technical measures as classified by TRAINS. These are the ones we drop from the SPS and TBT measures as they are the types of measures which are more likely to affect importers and not domestic firms. The NTM measures are aggregated to the 4 digit ISIC industry level. The total number of measures in each industry-year are summed and then divided by the number of HS6 products in the industry. Each row interacts the NTM measure with a dummy for quality, where quality is proxied by capital per worker, total wages per worker, and input expenditure per worker respectively. Firms are high-quality if the quality proxy is above the median in that industry in 1995. In all specifications we include an interaction of industry openness with the quality indicator, an interaction of the quality indicator measure with the industry import tariff, plus firm and industry-year interacted fixed effects. Standard errors – clustered by 4-digit industry – are in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

6.2 Model's Derivations

6.2.1 Consumers' Problem

Recall the Generalized Translated Power (GTP) preferences:

$$U = \int_{\Omega} \left(az\xi q(\omega) - \frac{(\xi q(\omega))^{1+\frac{1}{\gamma}}}{1+\frac{1}{\gamma}} \right) d\omega + \frac{\xi^{-\eta} - 1}{\eta} \tag{20}$$

where ξ is a quantity aggregator that is implicitly defined as:

$$\xi^{-\eta} = \int \left(az\xi q(\omega) - (\xi q(\omega))^{1+\frac{1}{\gamma}} \right) d\omega \tag{21}$$

The first order conditions of the consumers' problem are:

$$az\xi - \xi^{1+\frac{1}{\gamma}} q(\omega)^{\frac{1}{\gamma}} + \left[\underbrace{\int \left(azq(\omega) - \xi^{\frac{1}{\gamma}} (q(\omega))^{1+\frac{1}{\gamma}} \right) d\omega - \xi^{-\eta-1}}_{=0 \text{ by } (21)} \right] \frac{\partial \xi}{\partial q(\omega)} = \lambda p(\omega)$$

$$az\xi - \xi^{1+\frac{1}{\gamma}} q(\omega)^{\frac{1}{\gamma}} = \lambda p(\omega) \tag{22}$$

By multyplying both sides of (22) by $q(\omega)$ and integrating across all varieties $\omega \in \Omega$, we obtain the marginal utility of income λ.

$$\lambda = \frac{1}{y} \int \left(az\xi q(\omega) - (\xi q(\omega))^{1+\frac{1}{\gamma}} \right) d\omega = \frac{\xi^{-\eta}}{y} \tag{23}$$

Using (23) in (22) yields the inverse demand:

$$p(\omega) = \frac{\xi}{\lambda} \left[az(\omega) - (\xi q(\omega))^{\frac{1}{\gamma}} \right] = y\xi^{1+\eta} \left[az - (\xi q(\omega))^{\frac{1}{\gamma}} \right] \tag{24}$$

As we consider a closed economy, we normalize per capita income to unity.

6.2.2 Quality Standard and Aggregate Variables

The average profits of firms with $z > \bar{z}$ are:

$$\begin{aligned} \bar{\pi} &= \int_{\bar{z}}^{\infty} \pi(z) \frac{\kappa \bar{z}^{\kappa}}{z^{\kappa+1}} dz = \\ &= \frac{Lc}{1+\gamma} \left(\frac{a\gamma}{1+\gamma} \right)^{\gamma} \frac{(z^*)^{\gamma}}{\xi} \int_{\bar{z}}^{\infty} \left(\frac{z}{z^*} - 1 \right)^{1+\gamma} \frac{\kappa \bar{z}^{\kappa}}{z^{\kappa+1}} dz = \\ &= \frac{Lc}{1+\gamma} \left(\frac{a\gamma}{1+\gamma} \right)^{\gamma} \frac{(z^*)^{\gamma}}{\xi} G_1(g) \end{aligned} \tag{25}$$

where $G_1(g)$ is a function of κ, γ, and of the restrictiveness of the standard g:

$$\begin{aligned} G_1(g) &= \int_{\bar{z}}^{\infty} \kappa \left(\frac{z}{z^*} - 1\right)^{1+\gamma} \frac{\bar{z}^\kappa}{z^{\kappa+1}} dz = \\ &= \int_{\bar{z}}^{\infty} \kappa \left(1 - \frac{z^*}{z}\right)^{1+\gamma} \frac{\bar{z}^\kappa}{(z^*)^{1+\gamma} z^{\kappa-\gamma}} dz = \\ &= \kappa g^{1+\gamma} \left[\frac{F_1(g)}{\kappa - \gamma - 1} - g^{-1} \frac{F_2(g)}{\kappa - \gamma}\right] \end{aligned} \tag{26}$$

$F_1(g)$ and $F_2(g)$ are two hypergeometric functions given by:

$$F_1(g) = {}_2F_1\left[\kappa - \gamma - 1, -\gamma; \kappa - \gamma, g^{-1}\right]$$

$$F_2(g) = {}_2F_1\left[\kappa - \gamma, -\gamma; \kappa - \gamma + 1, g^{-1}\right].$$

The probability of a firm being active is:

$$P(z \geq \bar{z}) = \frac{b^\kappa}{\bar{z}^\kappa} = \frac{b^\kappa}{(z^* g)^\kappa} \tag{27}$$

The zero expected profit condition is:

$$\begin{aligned} P(z \geq \bar{z})\bar{\pi} &= f_E \\ \frac{Lc}{1+\gamma} \left(\frac{a\gamma}{1+\gamma}\right)^\gamma \frac{b^\kappa}{(z^*)^{\kappa-\gamma} g^\kappa \xi} G_1(g) &= f_E \end{aligned} \tag{28}$$

from which we obtain:

$$(z^*)^{\kappa-\gamma} \xi = \frac{Lcb^\kappa}{f_E(1+\gamma)} \left(\frac{a\gamma}{1+\gamma}\right)^\gamma g^{-\kappa} G_1(g) \tag{29}$$

Substituting the quality cutoff $z^* = \frac{c}{a}\xi^{-(1+\eta)}$ into (29) yields the quality cutoff z^* and market aggregator ξ as a function of g and model's parameters:

$$z^* = \left[\frac{Lc^{\frac{\eta}{1+\eta}} \gamma^\gamma b^\kappa a^{\gamma + \frac{1}{1+\eta}}}{f_E(1+\gamma)^{1+\gamma}} g^{-\kappa} G_1(g)\right]^{\frac{1}{\kappa - \gamma - \frac{1}{1+\eta}}} \tag{30}$$

$$\xi = \left[\frac{Lb^\kappa a^\kappa \gamma^\gamma}{f_E(1+\gamma)^{1+\gamma} c^{\kappa-\gamma-1}} g^{-\kappa} G_1(g)\right]^{-\frac{1}{(1+\eta)(\kappa-\gamma)-1}} \tag{31}$$

Firms' average revenues are:

$$\bar{r} = \int_{\bar{z}}^{\infty} r(z) \kappa \frac{\bar{z}^\kappa}{z^{\kappa+1}} dz = \frac{Lc}{1+\gamma} \left(\frac{a\gamma}{1+\gamma}\right)^\gamma \frac{(z^*)^\gamma}{\xi} G_2(g) \tag{32}$$

where $G_2(g)$ is a function of κ, γ, and g:

$$\begin{aligned} G_2(g) &= \int_{\bar{z}}^{\infty} \kappa \left(\frac{z}{z^*} - 1\right)^{\gamma} \left(\frac{z}{z^*} + \gamma\right) \frac{\bar{z}^{\kappa}}{z^{\kappa+1}} dz = \\ &= \int_{\bar{z}}^{\infty} \kappa \left(1 - \frac{z^*}{z}\right)^{\gamma} \left(1 + \gamma \frac{z^*}{z}\right) \frac{\bar{z}^{\kappa}}{(z^*)^{1+\gamma} z^{\kappa-\gamma}} dz = \\ &= \kappa g^{1+\gamma} \left[\frac{F_1(g)}{\kappa - \gamma - 1} + \gamma g^{-1} \frac{F_2(g)}{\kappa - \gamma}\right] \end{aligned} \tag{33}$$

Revenues normalized by average revenues, which we use in the calibration exercise, become:

$$\frac{r(z)}{\bar{r}} = (G_2(g))^{-1} \left(\frac{z}{z^*} - 1\right)^{\gamma} \left(\frac{z}{z^*} + \gamma\right) \tag{34}$$

By market clearing:

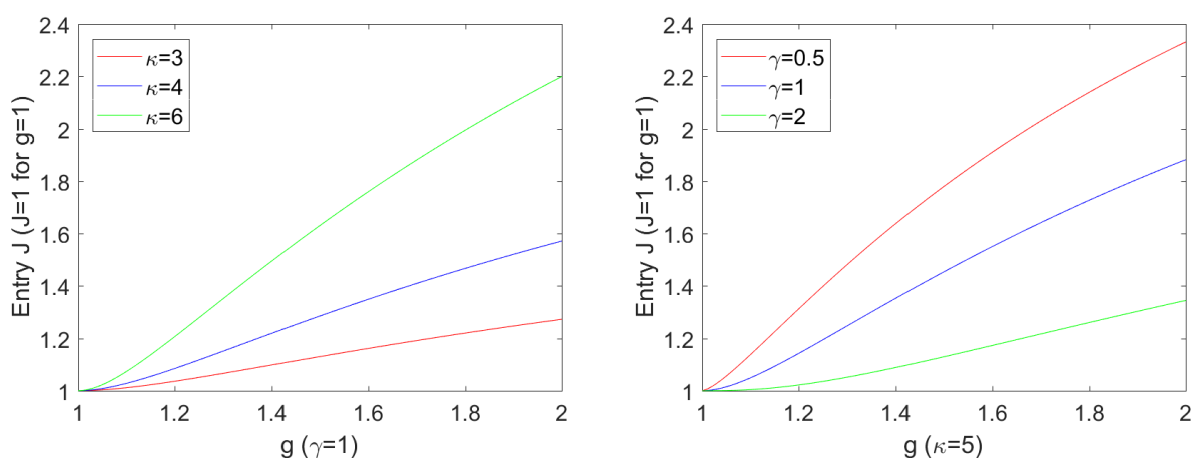
$$\frac{c}{1+\gamma} \left(\frac{a\gamma}{1+\gamma}\right)^{\gamma} \frac{Jb^{\kappa}}{(z^*)^{\kappa-\gamma} g^{\kappa} \xi} G_2(g) = 1 \tag{35}$$

Dividing (35) by (28) yields the equilibrium mass of entrants, which is independent of η:

$$J = \frac{L}{f_E} \frac{G_1(g)}{G_2(g)}$$

As shown in figure 5, market entry J is increasing in the restrictiveness of the standard.

Figure 5: Effects of a Standard on Entry



6.2.3 Welfare

To derive the utility, we need to derive the two integrals in (20) and (21). First, we obtain:

$$\int_{\bar{z}}^{\infty} a\xi z q(z) = a \left(\frac{a\gamma}{1+\gamma}\right)^{\gamma} \frac{Jb^{\kappa}}{(z^*)^{\kappa-\gamma} g^{\kappa}} z^* G_3(g) \tag{36}$$

where $G_3(g)$ is given by:

$$\begin{aligned} G_3(g) &= \int_{\bar{z}}^{\infty} \kappa \frac{z}{z^*} \left(\frac{z}{z^*} - 1 \right)^{\gamma} \frac{\bar{z}^{\kappa}}{z^{\kappa+1}} dz = \\ &= \int_{\bar{z}}^{\infty} \kappa \left(1 - \frac{z^*}{z} \right)^{\gamma} \frac{\bar{z}^{\kappa}}{(z^*)^{1+\gamma} z^{\kappa-\gamma}} dz = \\ &= \kappa g^{1+\gamma} \left[\frac{F_1(g)}{\kappa - \gamma - 1} \right] \end{aligned} \tag{37}$$

Rearranging the market clearing condition,

$$\left(\frac{a\gamma}{1+\gamma} \right)^{\gamma} \frac{Jb^{\kappa}}{(z^*)^{\kappa-\gamma} g^{\kappa} \xi} = \frac{(1+\gamma)\xi}{cG_2(g)} \tag{38}$$

Using (38) into (36) yields:

$$\int_{\bar{z}}^{\infty} a\xi z q(z) = (1+\gamma) \frac{az^*\xi}{c} \left(\frac{G_3(g)}{G_2(g)} \right) \tag{39}$$

Following the same steps, we obtain the second integral:

$$\begin{aligned} \int_{\bar{z}}^{\infty} (\xi q(z))^{1+\frac{1}{\gamma}} &= \left(\frac{a\gamma}{1+\gamma} \right)^{1+\gamma} \frac{Jb^{\kappa}}{(z^*)^{\kappa-\gamma} g^{\kappa}} z^* G_1(g) \\ &= \frac{az^*\xi}{c} \gamma \left(\frac{G_1(g)}{G_2(g)} \right) \end{aligned} \tag{40}$$

Substituting (39) and (40) into the utility function 20 yields:

$$\begin{aligned} U &= \int_{\Omega} \left(az\xi q(\omega) - \frac{(\xi q(\omega))^{1+\frac{1}{\gamma}}}{1+\frac{1}{\gamma}} \right) d\omega + \frac{\xi^{-\eta} - 1}{\eta} \\ &= \frac{az^*\xi}{c} \left[(1+\gamma) \frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma} \frac{G_1(g)}{G_2(g)} \right] + \frac{\xi^{-\eta} - 1}{\eta} \end{aligned}$$

By the cutoff condition (6), $\xi^{-\eta} = \frac{a\xi z^*}{c}$. Thus, the utility becomes:

$$U = \frac{az^*\xi}{c} \left[(1+\gamma) \frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma} \frac{G_1(g)}{G_2(g)} + \frac{1}{\eta} \right] - \frac{1}{\eta}$$

Finally, by the cutoff condition, $z^* = \frac{c}{a} \xi^{-1-\eta}$. Thus, the utility becomes:

$$U = \xi^{-\eta} \left[(1+\gamma) \frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma} \frac{G_1(g)}{G_2(g)} + \frac{1}{\eta} \right] - \frac{1}{\eta} \tag{41}$$

Substituting (31) into (41) yields:

$$U = \left[\frac{Lb^{\kappa}a^{\kappa}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}g^{-\kappa}G_1(g)\right]^{\frac{\eta}{(1+\eta)(\kappa-\gamma)-1}}\left[(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)} + \frac{1}{\eta}\right] - \frac{1}{\eta} \tag{42}$$

The government chooses $\bar{z}$. The equilibrium value of $z^*(g)$ is determined by the equation that describes the cutoff as a function of $z^*(g)$. The measure of the restrictiveness of the standard as the ratio between $\bar{z}$ and the market cutoff under no restriction $z^*(1)$ is given by:

$$\tilde{g} = \frac{\bar{z}}{z^*(1)} = \frac{\bar{z}}{z^*(g)}\frac{z^*(g)}{z^*(1)} = g\left[\frac{g^{-\kappa}G_1(g)}{G_1(1)}\right]^{\frac{1}{\kappa-\gamma-\frac{1}{1+\eta}}} \tag{43}$$

and exactly equals g under IA preferences $\eta = -1$.

Directly Additive Preferences

The case of DA preferences is obtained by setting $\eta \to \infty$. The utility, market cutoff, and aggregator become:

$$U_{DA} = \left[\frac{Lb^{\kappa}a^{\kappa}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}g^{-\kappa}G_1(g)\right]^{\frac{1}{\kappa-\gamma}}\left[(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}\right] \tag{44}$$

$$z^*_{DA}(g) = \left[\frac{Lc\gamma^{\gamma}b^{\kappa}a^{\gamma}}{f_E(1+\gamma)^{1+\gamma}}g^{-\kappa}G_1(g)\right]^{\frac{1}{\kappa-\gamma}} \tag{45}$$

$$\xi_{DA} = 1 \tag{46}$$

Indirectly Additive Preferences

The case of IA preferences is obtained by setting $\eta = -1$. The utility becomes:

$$U_{IA} = \left[\frac{Lb^{\kappa}a^{\kappa}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}g^{-\kappa}G_1(g)\right]\left[(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)} - 1\right] + 1 \tag{47}$$

$$z^*_{IA} = \frac{c}{a} \tag{48}$$

$$\xi_{IA} = \frac{Lb^{\kappa}a^{\kappa}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}g^{-\kappa}G_1(g) \tag{49}$$

Homothetic Preferences

The case of homothetic preferences is obtained by setting $\eta = 0$. The market determined cutoff and aggregator become:

$$z_H^* = \left[\frac{L\gamma^\gamma b^\kappa a^{\gamma+1}}{f_E(1+\gamma)^{1+\gamma}} g^{-\kappa} G_1(g)\right]^{\frac{1}{\kappa-\gamma-1}} \tag{50}$$

$$\xi_H = \left[\frac{Lb^\kappa a^\kappa \gamma^\gamma}{f_E(1+\gamma)^{1+\gamma}} g^{-\kappa} G_1(g) c^{\kappa-\gamma-1}\right]^{-\frac{1}{\kappa-\gamma-1}} \tag{51}$$

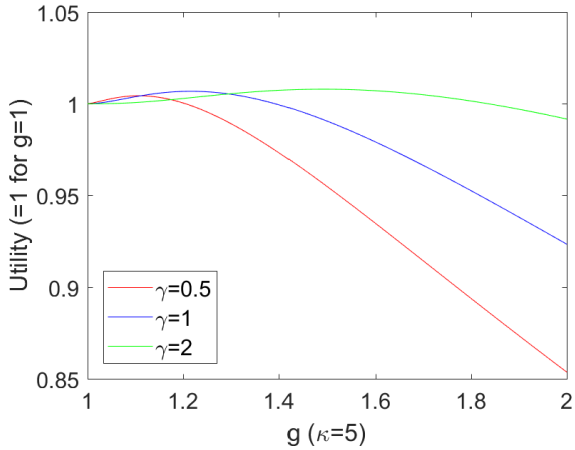
The utility becomes:

$$\begin{aligned} U_H &= \lim_{\eta\to 0}\left\{\left(\frac{az^*}{c}\right)^{\frac{\eta}{1+\eta}}\left[(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}\right] + \frac{\left(\frac{az^*}{c}\right)^{\frac{\eta}{1+\eta}} - 1}{\eta}\right\} = \\ &= \left[(1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}\right] + \ln\left(\frac{az^*}{c}\right) = \\ &= \frac{\ln\left[\frac{L\gamma^\gamma b^\kappa a^{\gamma+1}}{f_E(1+\gamma)^{1+\gamma}}\right]}{\kappa-\gamma-1} + \ln\left(\frac{a}{c}\right) + (1+\gamma)\frac{G_3(g)}{G_2(g)} - \frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)} - \frac{\kappa\ln g}{\kappa-\gamma-1} + \frac{\ln G_1(g)}{\kappa-\gamma-1} \end{aligned} \tag{52}$$

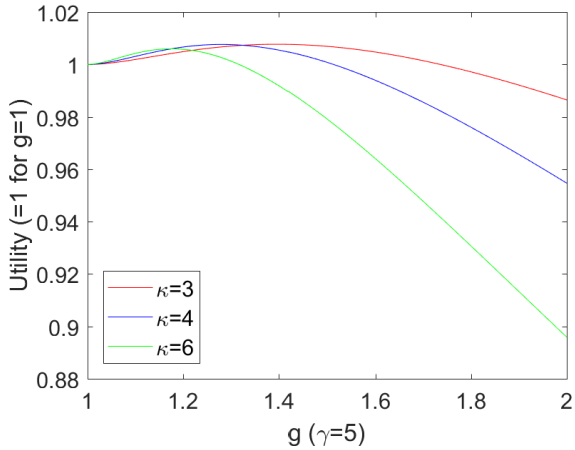
Figure 6 shows the hump-shaped relationship between the utility of the consumers and the restrictiveness of the standard g. Figure 7 presents the negative relationship between the mass of active firms N and the standard g. Figure 8 presents the negative relationship between the market determined quality cutoff z^* and the standard g, under DA and homothetic preferences, since under IA, the cutoff is a constant. Finally, figure 9 presents the relationship between the aggregator ξ and the standard g, under IA and homothetic preferences, since under DA, the aggregator equals one. The standard reduces the aggregator under IA preferences while it increases the aggregator under homothetic preferences.

Figure 6: Effects of a Standard on Welfare

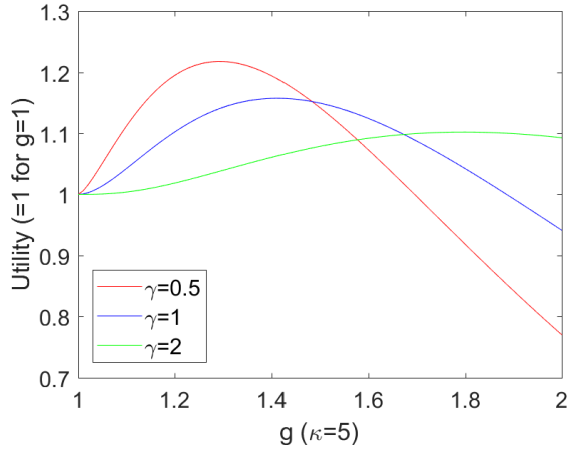
(a) Directly Additive



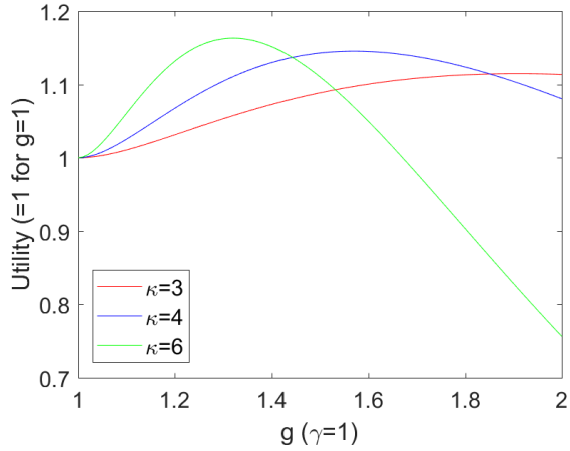
(b) Directly Additive



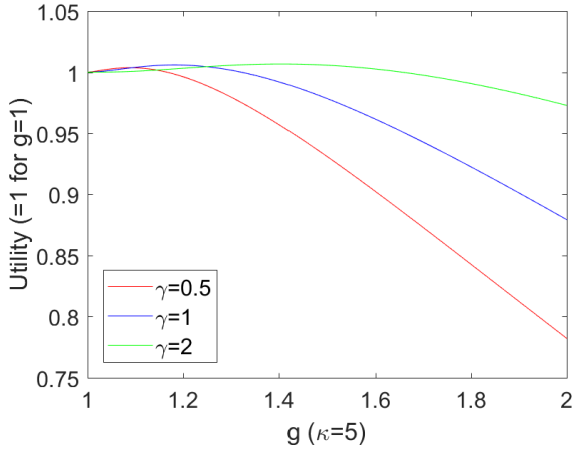
(c) Indirectly Additive



(d) Indirectly Additive



(e) Homothetic



(f) Homothetic

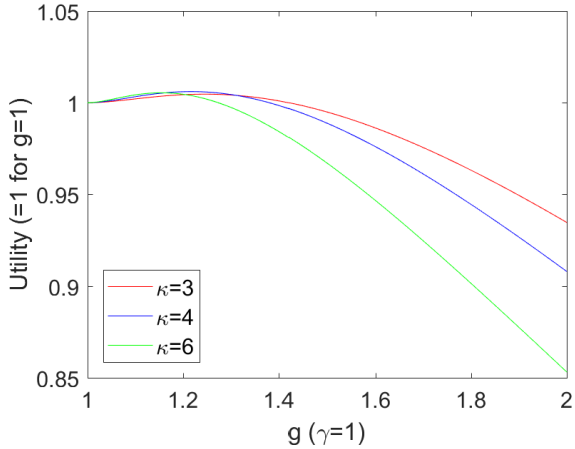
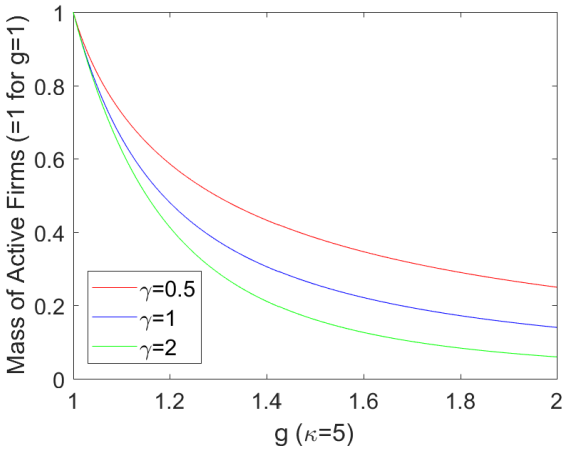
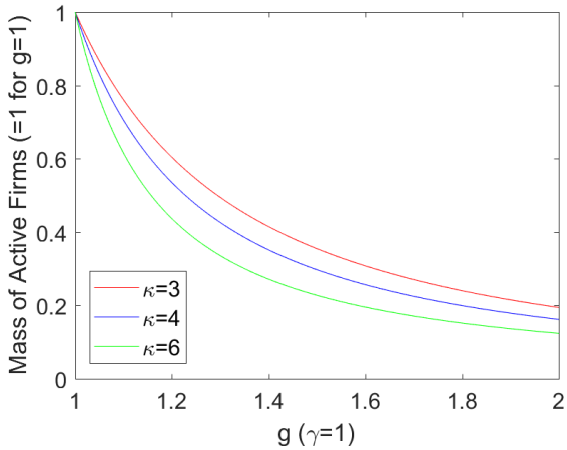


Figure 7: Effects of a Standard on the Mass of Active Firms N

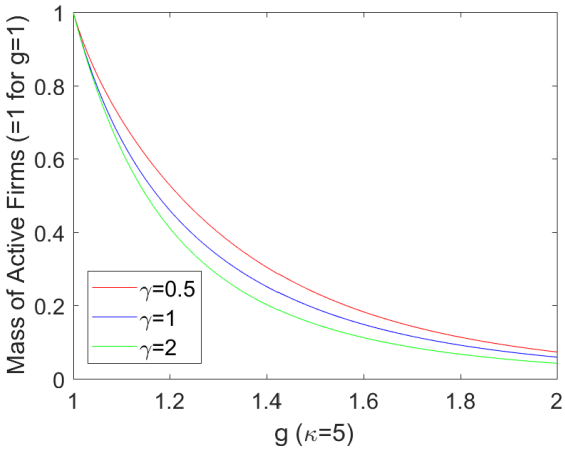
(a) Directly Additive



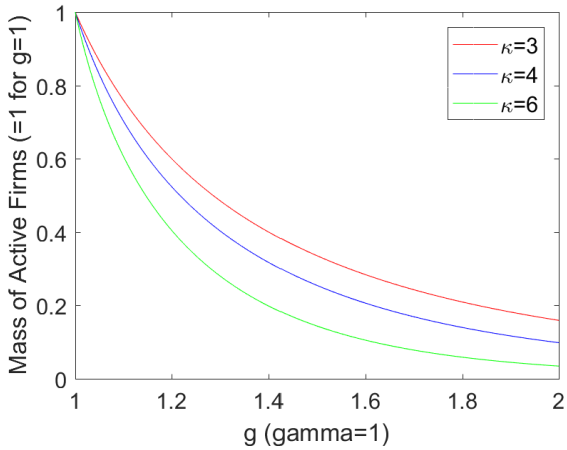
(b) Directly Additive



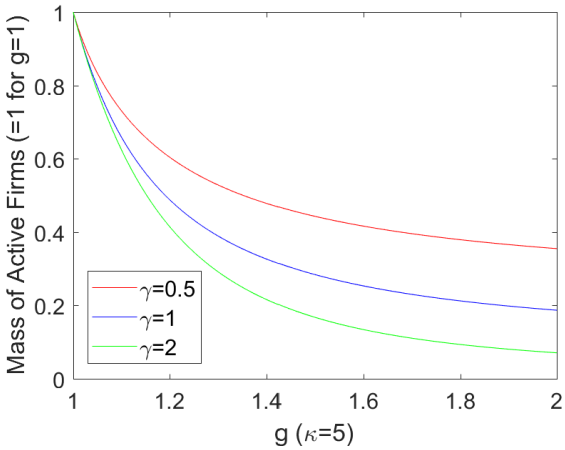
(c) Indirectly Additive



(d) Indirectly Additive



(e) Homothetic



(f) Homothetic

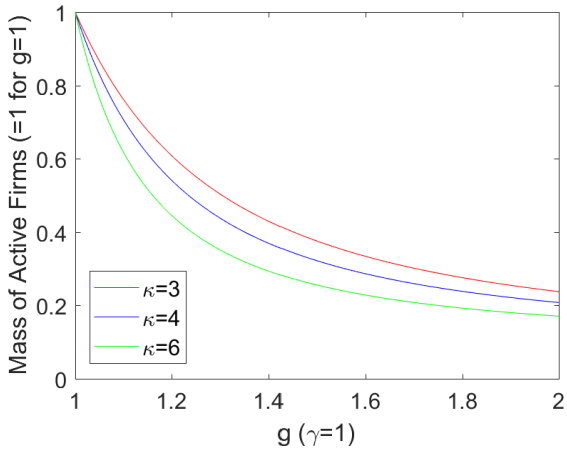
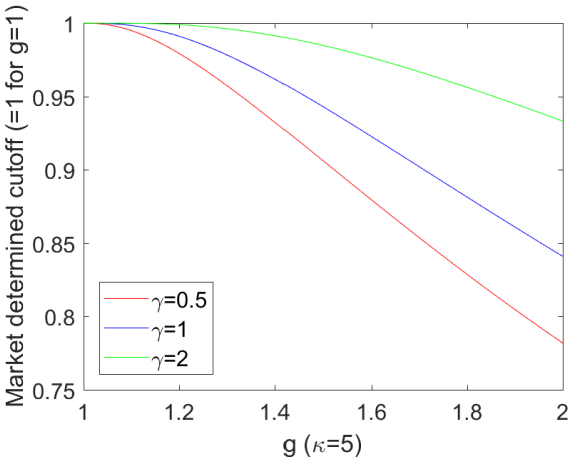
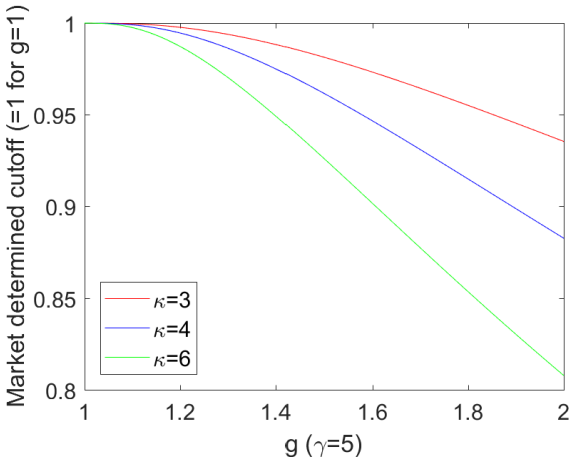


Figure 8: Effects of a Standard on the Market Quality Cutoff z^*

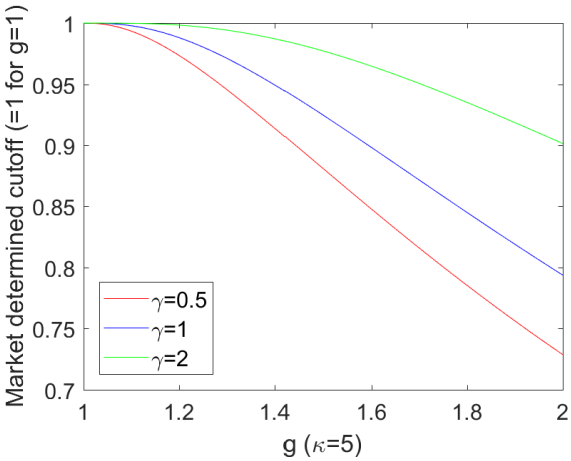
(a) Directly Additive



(b) Directly Additive



(c) Homothetic



(d) Homothetic

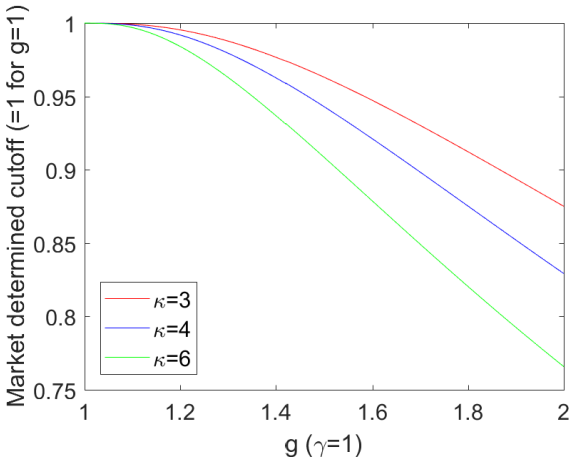
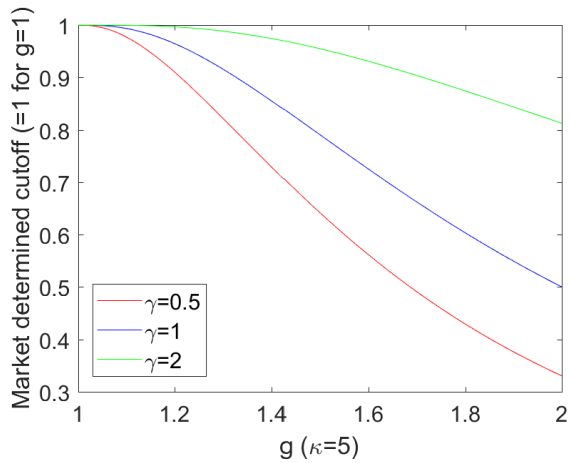
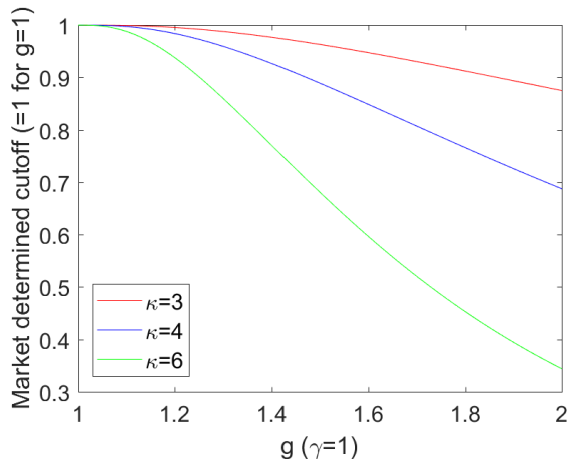


Figure 9: Effects of a Standard on the Aggregator ξ

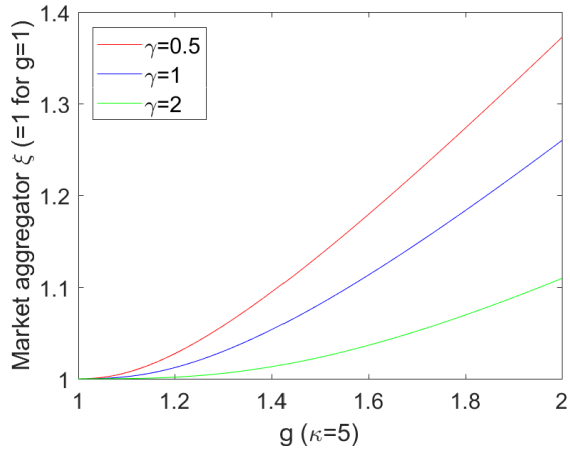
(a) Indirectly Additive



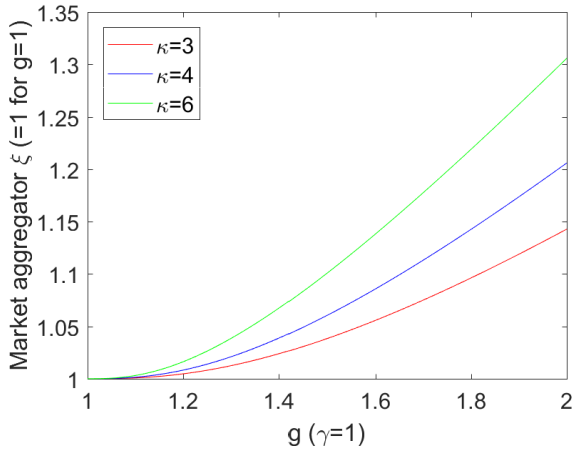
(b) Indirectly Additive



(c) Homothetic



(d) Homothetic



6.2.4 Fixed Cost

This section briefly outlines the case in which the government imposes a quality standard in the market $\bar{z}$, through a fixed cost of production f. The fixed cost f rationalizes the compliance costs that firms must incur due to the standard, or the costs associated with inspections for quality levels. The presence of a fixed cost f leaves the solution to the firms' problem (quantities and prices) and, thus, the revenues, unchanged. However, the profits of a firm with quality z become:

$$\pi(z) = \frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z^*)^{\gamma}}{\xi}\left(\frac{z}{z^*}-1\right)^{1+\gamma} - f$$

There exists a firm with quality $\bar{z}$, such that $\pi(\bar{z}) = 0$. The mapping between the fixed cost f and the cutoff-firm $\bar{z}$, equivalent to the quality standard imposed in the baseline model,

is:

$$\frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z^*)^{\gamma}}{\xi}\left(\frac{\bar{z}}{z^*}-1\right)^{1+\gamma}=f$$
$$\frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z^*)^{\gamma}}{\xi}(g-1)^{1+\gamma}=f \tag{53}$$

Using (53), average profits become:

$$\begin{aligned}\bar{\pi}&=\frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z^*)^{\gamma}}{\xi}G_1(g)-f\\&=\frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z^*)^{\gamma}}{\xi}(G_1(g)-(g-1)^{1+\gamma})\end{aligned}$$

The zero expected profit condition is then

$$\frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{b^{\kappa}}{(z^*)^{\kappa-\gamma}g^{\kappa}\xi}(G_1(g)-(g-1)^{1+\gamma})=f_E$$

Relative to the baseline model, the fixed cost affects both z^* and ξ. Using the market quality cutoff condition $z^*=\frac{c}{a}\xi^{-(1+\eta)}$ into the zero expected profit condition yields the solutions for z^* and ξ. The fixed cost affect the utility of the representative consumer (41) only through ξ. Following the same steps of the baseline model, the utility becomes:

$$U=\left[\frac{Lb^{\kappa}a^{\kappa}c^{\kappa-\gamma-1}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}}g^{-\kappa}(G_1(g)-(g-1)^{1+\gamma})\right]^{\frac{\eta}{(1+\eta)(\kappa-\gamma)-1}}\left[(1+\gamma)\frac{G_3(g)}{G_2(g)}-\frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}+\frac{1}{\eta}\right]-\frac{1}{\eta}$$

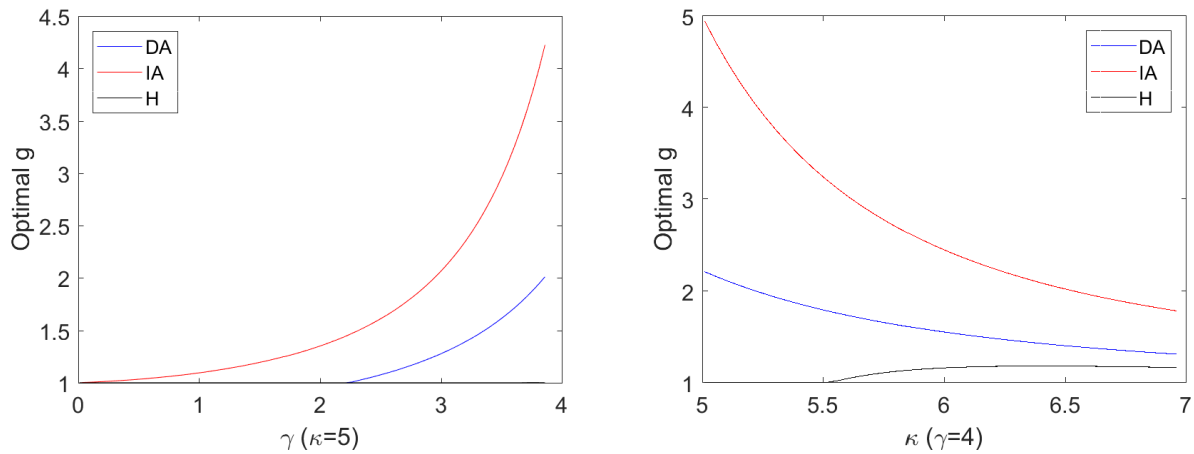
and deriving the closed form expressions for the IA, DA, and homothetic case follows from the baseline model. In particular,

$$\begin{aligned}U_{DA}&=\left[\frac{Lb^{\kappa}a^{\kappa}c^{\kappa-\gamma-1}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}}g^{-\kappa}(G_1(g)-(g-1)^{1+\gamma})\right]^{\frac{1}{\kappa-\gamma}}\left[(1+\gamma)\frac{G_3(g)}{G_2(g)}-\frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}\right]\\U_{IA}&=\left[\frac{Lb^{\kappa}a^{\kappa}c^{\kappa-\gamma-1}\gamma^{\gamma}}{f_E(1+\gamma)^{1+\gamma}}g^{-\kappa}(G_1(g)-(g-1)^{1+\gamma})\right]\left[(1+\gamma)\frac{G_3(g)}{G_2(g)}-\frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}-1\right]+1\\U_H&=\frac{\ln\left[\frac{L\gamma^{\gamma}b^{\kappa}a^{\gamma+1}}{f_E(1+\gamma)^{1+\gamma}}\right]}{\kappa-\gamma-1}+\ln\left(\frac{a}{c}\right)+(1+\gamma)\frac{G_3(g)}{G_2(g)}-\frac{\gamma^2}{1+\gamma}\frac{G_1(g)}{G_2(g)}-\frac{\kappa\ln g}{\kappa-\gamma-1}+\frac{\ln(G_1(g)-(g-1)^{1+\gamma})}{\kappa-\gamma-1}\end{aligned}$$

While the standard in the baseline model directly affect the selection of firms, the fixed cost also reallocates resources from production to the activities required to comply to the regulation. As a result, the welfare benefits of the standard examined in the baseline case are diminished by the fixed cost. In fact, the optimal standard is smaller in this extension that it is in the baseline case: the fixed cost acts as a downward shift in the optimal g across preferences. As shown in figure 10, for the DA and homothetic case the optimal policy is no standard under certain parameters.

Figure 10: Optimal g Across Preferences - Standard as Fixed Cost

(a) As a function of γ **(b)** As a function of κ



6.2.5 Productivity Heterogeneity and Quality

The baseline model features the simplifying assumption that firms differ exogenously in terms of quality. However, most papers in the literature model firms that differ in terms of productivity and that product quality is a function of firm's productivity (Manova and Zhang, 2017; Feenstra and Weinstein, 2016). This section shows that the results of our baseline model also arise in a model in which quality is a function of firm's productivity.

Consider an extension to the baseline model in which firms differ in terms of productivity ϕ. As it is common in the literature, we assume that ϕ follows a Pareto distribution with CDF: $1 - \left(\frac{\tilde{b}}{\phi}\right)^{\tilde{\kappa}}$. Similarly to the framework of Manova and Zhang (2017), firm's quality is proportional to firm's productivity: $z = \phi^{\frac{1}{\theta}}$, with $\theta > 0$. Moreover, we let the marginal cost of the firm ϕ be proportional to the quality. In particular, marginal costs are equal to cz^{β}. We assume that the elasticity of marginal costs with respect to quality is less than one: $\beta < 1$. This assumption is made for average revenues to be well defined[57]. To obtain a closed form expression for the utility, we restrict the analysis to the linear GTP case, namely $\gamma = 1$.

Firm's profits become

$$\pi(z) = L\xi^{1+\eta}\left[azq(z) - \xi(q(z))^2\right] - Lcz^{\beta}q(\omega)$$

Profit maximization yields the following optimal quantity:

$$q(z) = \left(\frac{a}{2}\right)\frac{z^*}{\xi}\left(\frac{z}{z^*} - \left(\frac{z}{z^*}\right)^{\beta}\right)$$

[57]Modeling an endogenous quality choice as (Gaigne and Larue, 2016), in which firms must also pay a fixed cost is highly untractable under GTP preferences. We verified that such a technological assumption does not generate additional distortions in a model with standard CES preferences.

where the market determined quality cutoff equals:

$$z^* = \left(\frac{c}{a\xi^{1+\eta}}\right)^{\frac{1}{1-\beta}}$$

Using the quality cutoff, we can rewrite the performance variables of the firm as follows:

$$\begin{aligned}
p(z) &= \frac{c(z^*)^\beta}{2}\left(\frac{z}{z^*} + \left(\frac{z}{z^*}\right)^\beta\right) \\
r(z) &= \frac{Lca}{4}\frac{(z^*)^{\beta+1}}{\xi}\left(\left(\frac{z}{z^*}\right)^2 - \left(\frac{z}{z^*}\right)^{2\beta}\right) \\
\pi(z) &= \frac{Lca}{4}\frac{(z^*)^{\beta+1}}{\xi}\left(\frac{z}{z^*} - \left(\frac{z}{z^*}\right)^\beta\right)^2
\end{aligned}$$

Let us derive the probability distribution for quality. In particular,

$$Pr(\tilde{z} \leq z) = Pr(\phi^{\frac{1}{\theta}} \leq z) = 1 - \left(\frac{\tilde{b}}{z^\theta}\right)^{\tilde{\kappa}}$$

Thus, we can change the notation and derive the same distribution for quality we used in the baseline model. In fact, quality z follows a Pareto distribution with shape parameter $\kappa = \tilde{\kappa}\theta$ and shift parameter $b = \tilde{b}^{\frac{1}{\theta}}$.

Following the same procedure as the baseline model, the utility of the representative consumer becomes:

$$U = \left[\frac{Lb^\kappa a^{\kappa-2\beta}}{4f_E c^{\kappa-2}} g^{-\kappa} G_1(g)\right]^{\frac{\eta(1-\beta)}{(1+\eta)(\kappa-1)-1-\beta\eta}} \left[2\frac{G_3(g)}{G_2(g)} - \frac{1}{2}\frac{G_1(g)}{G_2(g)} + \frac{1}{\eta}\right] - \frac{1}{\eta}$$

where

$$\begin{aligned}
G_1(g) &= \frac{\kappa g^2}{\kappa-2} - \frac{2\kappa g^{\beta+1}}{\kappa-\beta-1} + \frac{\kappa g^{2\beta}}{\kappa-2\beta} \\
G_2(g) &= \frac{\kappa g^2}{\kappa-2} - \frac{\kappa g^{2\beta}}{\kappa-2\beta} \\
G_3(g) &= \frac{\kappa g^2}{\kappa-2} - \frac{\kappa g^{\beta+1}}{\kappa-\beta-1}
\end{aligned}$$

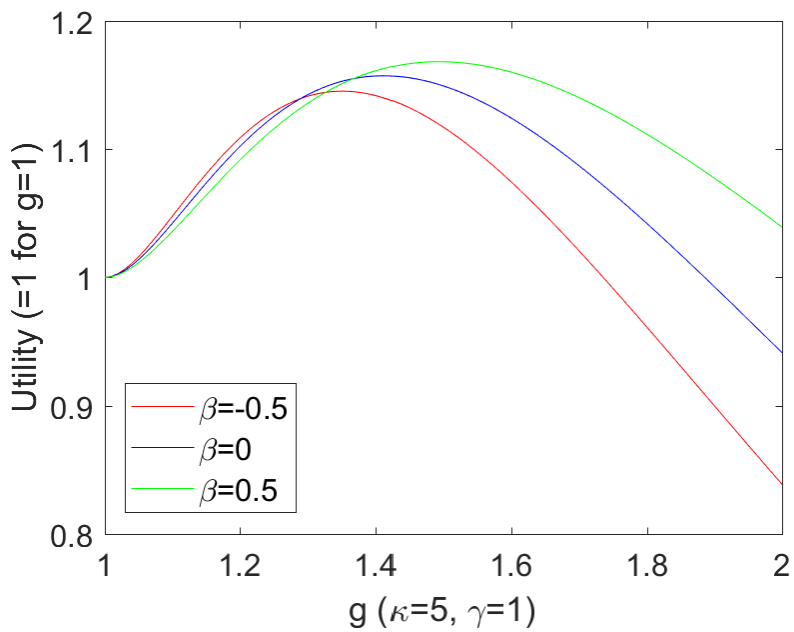
Figure 11 shows the relationship between welfare and the restrictiveness of the standard for different values of β, under IA preferences. For $\beta = 0$, this extension becomes identical to the baseline model. This implies that our baseline model with firms heterogeneous in quality is equivalent to a model in which firms differ in terms of productivity, and their productivity is proportional to their product quality. The result is independent of the level of θ, as long as the two models match the same distribution of sales.

For $\beta \neq 0$, the marginal costs of production depends on quality. If $\beta < 0$, firm's with high quality also have lower production costs. This scenario assumes that more productive

firms have higher quality and attain cost efficiency. When $\beta < 0$, the sales difference between low-quality firms and high-quality firms increases relative to the baseline model, since high-quality firms are also low-cost firms. In this case, the business stealing bias is reduced relative to the baseline model. Hence, the optimal level of g is smaller.

On the other hand, $\beta > 0$ yields the more realistic scenario in which high-quality firms have higher costs of production than low-quality firms (Manova and Zhang, 2017). In this scenario, the business stealing bias is larger than the baseline case. There are too many low-quality firms operating in the market because 1) their markups are lower and 2) their marginal costs are lower. As a result, when marginal costs and quality are positively correlated the positive welfare effects of the standard are larger. The result also arises under DA and homothetic preferences. Details are available upon request.

Figure 11: Minimum Quality Standard and Welfare ($\eta = -1$)



6.2.6 A Two-Country Framework

Consider a world made of two identical economies, where exporting requires an iceberg trade cost τ. Consumers have GTP preferences. As the two countries are identical, we can normalize per capita income and wages to one in both countries. We assume that the two countries decide a common minimum quality requirement $\bar{z}$, which holds in each nation, and we study how different levels of the iceberg trade cost τ affect the optimal level of the standard.

Let subscript D denote variables related to the domestic economy (domestic sales, profits and so on), and X variables related to exports. Adapting the market-determined quality cutoff to this section's assumptions yields the following market-determined cutoffs for domestic

sales z_D^* and for exports z_X^*:

$$z_D^* = \frac{c}{a}\xi^{1+\eta} \tag{54}$$

$$z_X^* = \frac{\tau c}{a}\xi^{1+\eta} \tag{55}$$

When the international authority chooses $\bar{z}$ there are two potential outcomes. First, if $\bar{z} \in [z_D^*, z_X^*]$, only domestic sales are affected: the low-quality firms that sell domestically are forced out of the market but the selection of exporters only depends on the iceberg trade cost τ. Second, if $\bar{z} \in [z_X^*, \infty)$, both domestic sellers and exporters are constrained by the policy. In particular, the minimum quality requirement is so demanding that only high-quality firms that exports are able to remain active. As a result, in this scenario — which appears unrealistic — all firms that are active also export. In this section, we focus on the first case, in which $\bar{z} \in [z_D^*, z_X^*]$. Details on the second case are available upon request.

Firm z profits and revenues in each market $i = D, X$ are identical to baseline model, with the exception that the cutoff z_i^* is destination specific:

$$r_i(z) = \frac{L\tau_i c}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z_i^*)^{\gamma}}{\xi}\left(\frac{z}{z_i^*}-1\right)^{\gamma}\left(\frac{z}{z_i^*}+\gamma\right)$$

$$\pi_i(z) = \frac{L\tau_i c}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z_i^*)^{\gamma}}{\xi}\left(\frac{z}{z_i^*}-1\right)^{1+\gamma}$$

where $\tau_D = 1$ and $\tau_X = \tau$. Let $g = \frac{\bar{z}}{z_D^*}$. Average revenues in the domestic and foreign economy are given by:

$$\bar{r}_D = \int_{\bar{z}}^{\infty} r_D(z)\kappa\frac{\bar{z}^{\kappa}}{z^{\kappa+1}}dz = \frac{Lc}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z_D^*)^{\gamma}}{\xi}G_2(g)$$

$$\bar{r}_X = \int_{z_X^*}^{\infty} r_X(z)\kappa\frac{(z^*)_X^{\kappa}}{z^{\kappa+1}}dz = \frac{L\tau c}{1+\gamma}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\frac{(z_X^*)^{\gamma}}{\xi}G_2(1)$$

The mass of active domestic firms is $J\left(\frac{b^{\kappa}}{\bar{z}^{\kappa}}\right)$ while the mass of exporters is $J\left(\frac{b^{\kappa}}{{z_X^*}^{\kappa}}\right)$. By (54) and (55), $z_X^* = \tau z_D^*$. Therefore, aggregate revenues R are given by:

$$\begin{aligned}
R &= J\left(\frac{b^{\kappa}}{\bar{z}^{\kappa}}\right)\bar{r}_D + J\left(\frac{b^{\kappa}}{(z^*)_X^{\kappa}}\right)\bar{r}_X = \frac{Jb^{\kappa}Lc}{(1+\gamma)\xi}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\left[\frac{G_2(g)(z_D^*)^{\gamma}}{\bar{z}^{\kappa}} + \frac{\tau G_2(1)}{(z_X^*)^{\kappa-\gamma}}\right] \\
&= \frac{Jb^{\kappa}Lc}{(1+\gamma)\xi}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\left[\frac{G_2(g)(z_D^*)^{\gamma}}{\bar{z}^{\kappa}} + \frac{\tau G_2(1)}{(z_X^*)^{\kappa-\gamma}}\right] = \\
&= \frac{Jb^{\kappa}Lc}{(1+\gamma)\xi(z_D^*)^{\kappa-\gamma}}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\left[\frac{G_2(g)}{g^{\kappa}} + \frac{G_2(1)}{\tau^{\kappa-\gamma-1}}\right]
\end{aligned}$$

Our market clearing condition ($R = L$) becomes:

$$\frac{Jb^{\kappa}c}{(1+\gamma)\xi(z_D^*)^{\kappa-\gamma}}\left(\frac{a\gamma}{1+\gamma}\right)^{\gamma}\left[g^{-\kappa}G_2(g) + \tau^{-(\kappa-\gamma-1)}G_2(1)\right] = 1 \tag{56}$$

Expected profits are computed as:

$$E[\pi] = \left(\frac{b^\kappa}{\bar{z}^\kappa}\right)\int_{\bar{z}}^{\infty} \pi_D(z)\kappa\frac{\bar{z}^\kappa}{z^{\kappa+1}}dz + \left(\frac{b^\kappa}{(z^*)_X^\kappa}\right)\int_{z_X^*}^{\infty} \pi_X(z)\kappa\frac{(z^*)_X^\kappa}{z^{\kappa+1}}dz$$

The zero expected profit condition ($E[\pi] = f_E$) becomes:

$$\frac{Lb^\kappa c}{(1+\gamma)\xi(z_D^*)^{\kappa-\gamma}}\left(\frac{a\gamma}{1+\gamma}\right)^\gamma \left[g^{-\kappa}G_1(g) + \tau^{-(\kappa-\gamma-1)}G_1(1)\right] = f_E \tag{57}$$

Let $G_5(g,\tau) = g^{-\kappa}G_1(g) + \tau^{-(\kappa-\gamma-1)}G_1(1)$. Using (54), the aggregator is:

$$\xi = \left[\frac{Lb^\kappa a^\kappa \gamma^\gamma}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}G_5(g,\tau)\right]^{-\frac{1}{(1+\eta)(\kappa-\gamma)-1}} \tag{58}$$

Using the market clearing condition (56), the first component of the utility function is:

$$\begin{aligned}\int a\xi z q(z)dz &= \int_{\bar{z}}^{\infty} a\xi z q_D(z)dz + \int_{z_X^*}^{\infty} a\xi z q_X(z)dz = \\ &= a\left(\frac{a\gamma}{1+\gamma}\right)^\gamma \frac{Jb^\kappa}{(z_D^*)^{\kappa-\gamma-1}}(g^{-\kappa}G_3(g) + \tau^{-(\kappa-\gamma-1)}G_3(1)) = \\ &= \frac{a\xi z_D^*}{c}(1+\gamma)\frac{g^{-\kappa}G_3(g) + \tau^{-(\kappa-\gamma-1)}G_3(1)}{g^{-\kappa}G_2(g) + \tau^{-(\kappa-\gamma-1)}G_2(1)}\end{aligned}$$

Similarly, the second component becomes:

$$\int_{\bar{z}}^{\infty} (\xi q(z))^{1+\frac{1}{\gamma}} = \frac{a\xi z_D^*}{c}\gamma\frac{g^{-\kappa}G_1(g) + \tau^{-(\kappa-\gamma-1)}G_1(1)}{g^{-\kappa}G_2(g) + \tau^{-(\kappa-\gamma-1)}G_2(1)}$$

Let $G_6(g,\tau)$ be equal to:

$$G_6(g,\tau) = (1+\gamma)\frac{g^{-\kappa}G_3(g) + \tau^{-(\kappa-\gamma-1)}G_3(1)}{g^{-\kappa}G_2(g) + \tau^{-(\kappa-\gamma-1)}G_2(1)} - \frac{\gamma^2}{1+\gamma}\frac{g^{-\kappa}G_1(g) + \tau^{-(\kappa-\gamma-1)}G_1(1)}{g^{-\kappa}G_2(g) + \tau^{-(\kappa-\gamma-1)}G_2(1)}$$

Using the previous results, and the definition of the cutoff (54), the utility equals to:

$$\begin{aligned}U &= \frac{a\xi z_D^*}{c}G_6(g,\tau) + \frac{\xi^{-\eta}-1}{\eta} = \\ &= \xi^{-\eta}\left[G_6(g,\tau) + \frac{1}{\eta}\right] - \frac{1}{\eta} = \\ &= \left[\frac{Lb^\kappa a^\kappa \gamma^\gamma}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}G_5(g,\tau)\right]^{\frac{\eta}{(1+\eta)(\kappa-\gamma)-1}}\left[G_6(g,\tau) + \frac{1}{\eta}\right] - \frac{1}{\eta}\end{aligned}$$

Let $\tilde{\xi} = \frac{Lb^\kappa a^\kappa \gamma^\gamma}{f_E(1+\gamma)^{1+\gamma}c^{\kappa-\gamma-1}}$. The utility under the three preferences (IA, DA, and homothetic)

is given by:

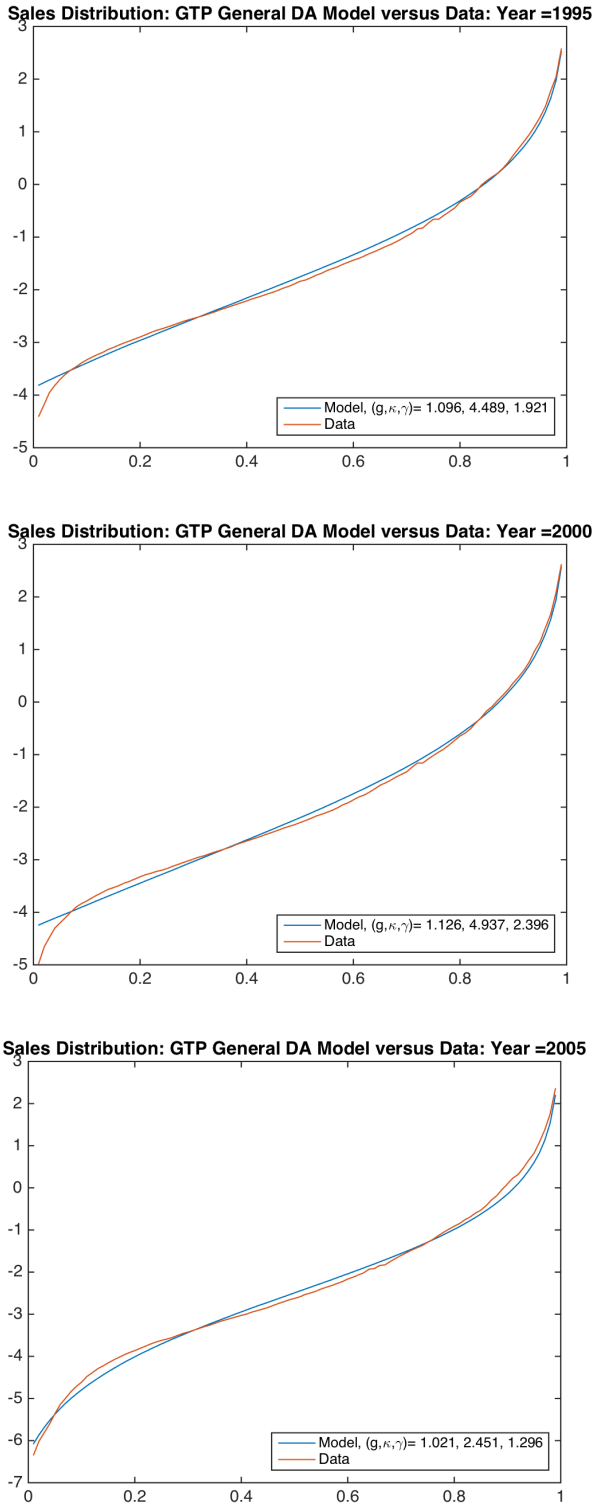
$$
\begin{aligned}
U_{IA} &= \tilde{\xi} G_5(g,\tau)\left[G_6(g,\tau)-1\right]+1 \\
U_{DA} &= (\tilde{\xi} G_5(g,\tau))^{\frac{1}{\kappa-\gamma}} G_6(g,\tau) \\
U_H &= \frac{\ln \tilde{\xi}}{\kappa-\gamma-1} + G_6(g,\tau) + \frac{\ln G_5(g,\tau)}{\kappa-\gamma-1}
\end{aligned}
$$

Under the three preferences, we compare the optimal g as a function of τ, restricting the values of τ such that $g^{opt} \in [1,\tau]$.

6.3 Estimation

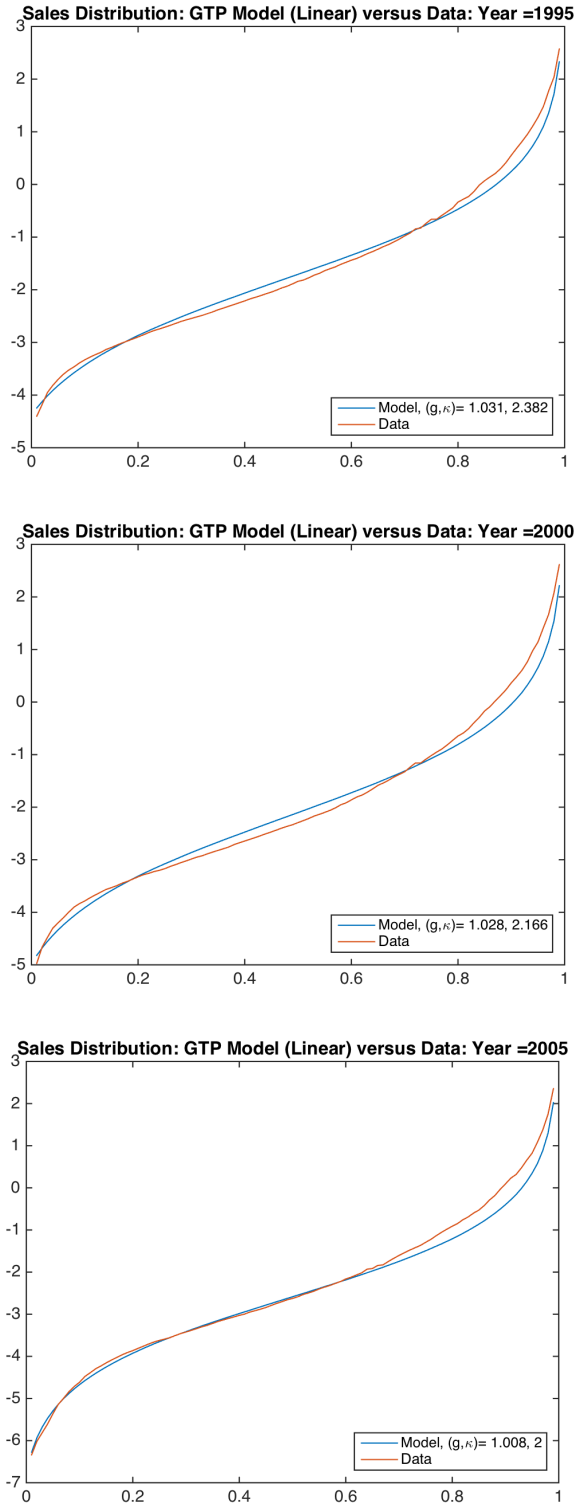
Figure 12 displays the results for the benchmark manufacturing-wide estimation for each year. We show results for the years 1995, 2000, 2005. Figure 13 displays the result under the alternative calibration in which we target only 4 moments: the sales advantage of "high-quality" relative to "low-quality" firms; the skewness of the distribution; and two differences: $\log(\tilde{r})_{99} - \log(\tilde{r})_{90}$ and $\log(\tilde{r})_{90} - \log(\tilde{r})_{10}$. Figure 14 plots the simulated sales distribution with a fixed linear demand: $\gamma = 1$. Figure 15 plots the simulated sales distribution when we set $\kappa = 4$ and $\gamma = 1.8$.

Figure 12: Log Domestic Sales Distribution: Model VS Data



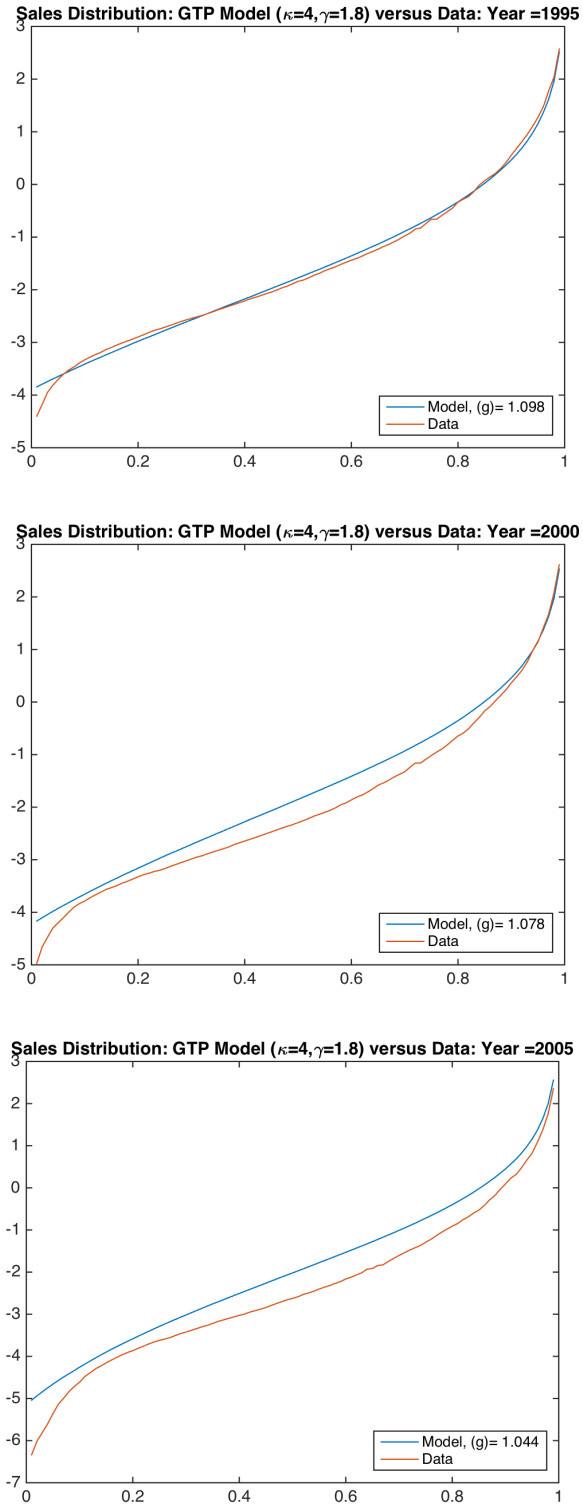
Each figure plots the CDF of the log sales distribution in the data (red) versus the simulated distribution given the estimated parameters. The model is estimated using the universe of manufacturing firms in each year. Although we estimate parameters for all years, we report only 1995, 2000, and 2005. The 95% confidence intervals for the parameters in each of the three years are: $\hat{g} = (1.08, 1.11), (1.1, 1.15), (1.01, 1.03)$; $\hat{\kappa} = (3.37, 5.6), (3.59, 6.29), (2.12, 2.78)$; $\hat{\gamma} = (1.63, 2.21), (2.04, 2.75), (1.17, 1.42)$.

Figure 14: Log Domestic Sales Distribution: Model ($\gamma = 1$) VS Data



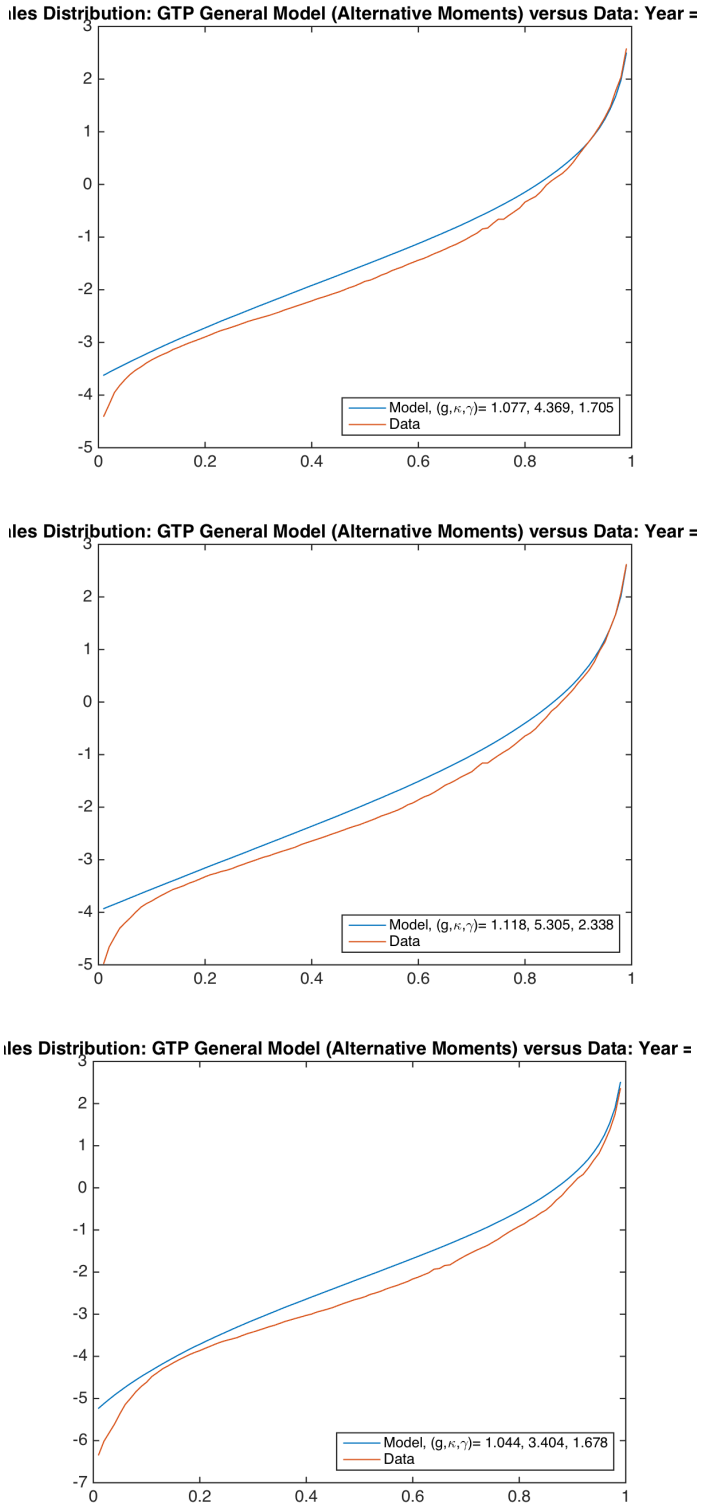
Each figure plots the CDF of the log sales distribution in the data (red) versus the simulated distribution given the estimated parameters. The parameters of the model are estimated using the main specification, but a fixed $\gamma = 1$. The model is estimated using the universe of manufacturing firms in each year. The 95% confidence intervals for the parameters in each of the three years are: $\hat{g} = (1.02, 1.04), (1.02, 1.03), (1.00, 1.02)$; $\hat{\kappa} = (0.94, 2.51), (1.26, 2.29), (2, 2)$.

Figure 15: Log Domestic Sales Distribution: Model ($\gamma = 1.8, \kappa = 4$) VS Data



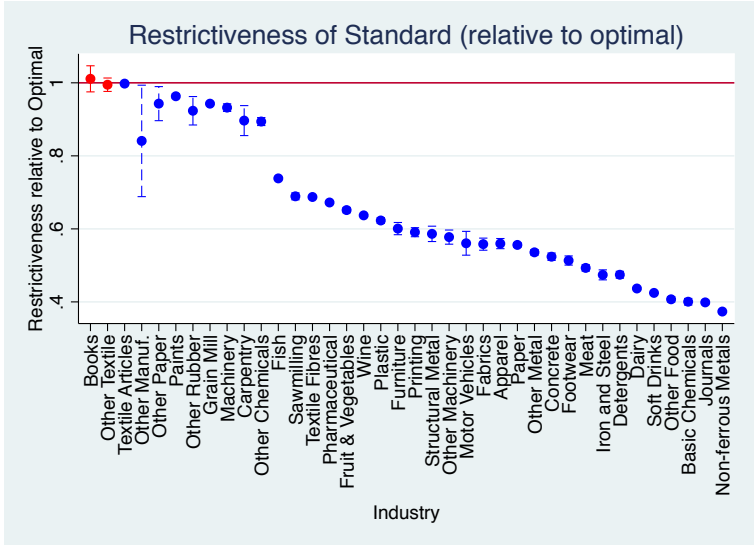
Each figure plots the CDF of the log sales distribution in the data (red) versus the simulated distribution given the estimated parameters. The parameters of the model are estimated using the main specification, but a fixed $\gamma = 1.8$ and $\kappa = 4$. The model is estimated using the universe of manufacturing firms in each year. The 95% confidence intervals for the parameters in each of the three years are: $\hat{g} = (1.09, 1.1), (1.07, 1.08), (1.04, 1.05)$.

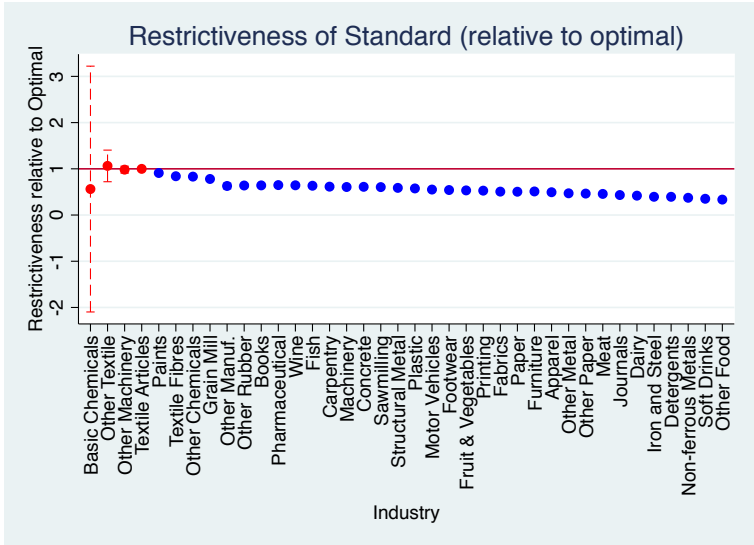
Figure 13: Log Domestic Sales Distribution: Model (Alternative Calibration) VS Data

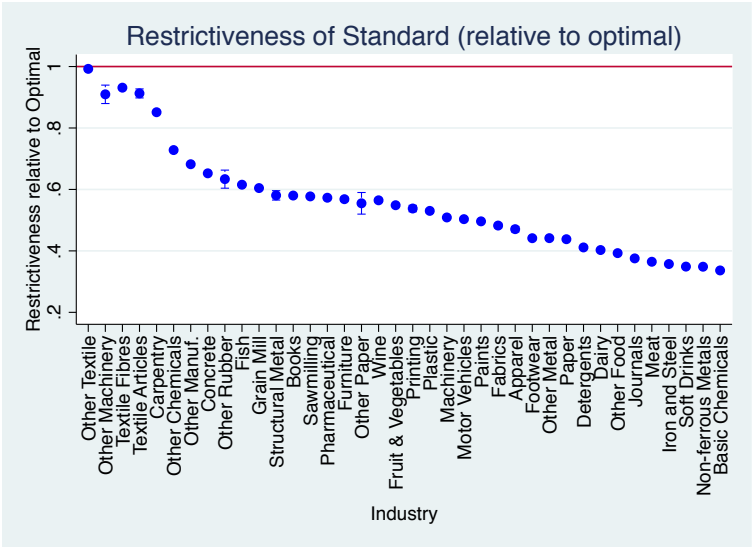


Each Figure plots the CDF of the log sales distribution in the data (red) versus the simulated distribution given the estimated parameters. The parameters of the model are estimated using *only the 4 moments described in the "alternative calibration"*. The model is estimated using the universe of manufacturing firms in each year. The 95% confidence intervals for the parameters in each of the three years are: $\hat{g} = (1.05, 1.11), (1.08, 1.14), (1.03, 1.06)$; $\hat{\kappa} = (0.94, 8.23), (1.26, 8), (2.08, 4.6)$; $\hat{\gamma} = (1.11, 2.50), (1.42, 2.87), (1.27, 2.06)$.

Figure 16: Restrictiveness Index in IA with $\gamma = 1$ Model: 1995, 2000, and 2005.

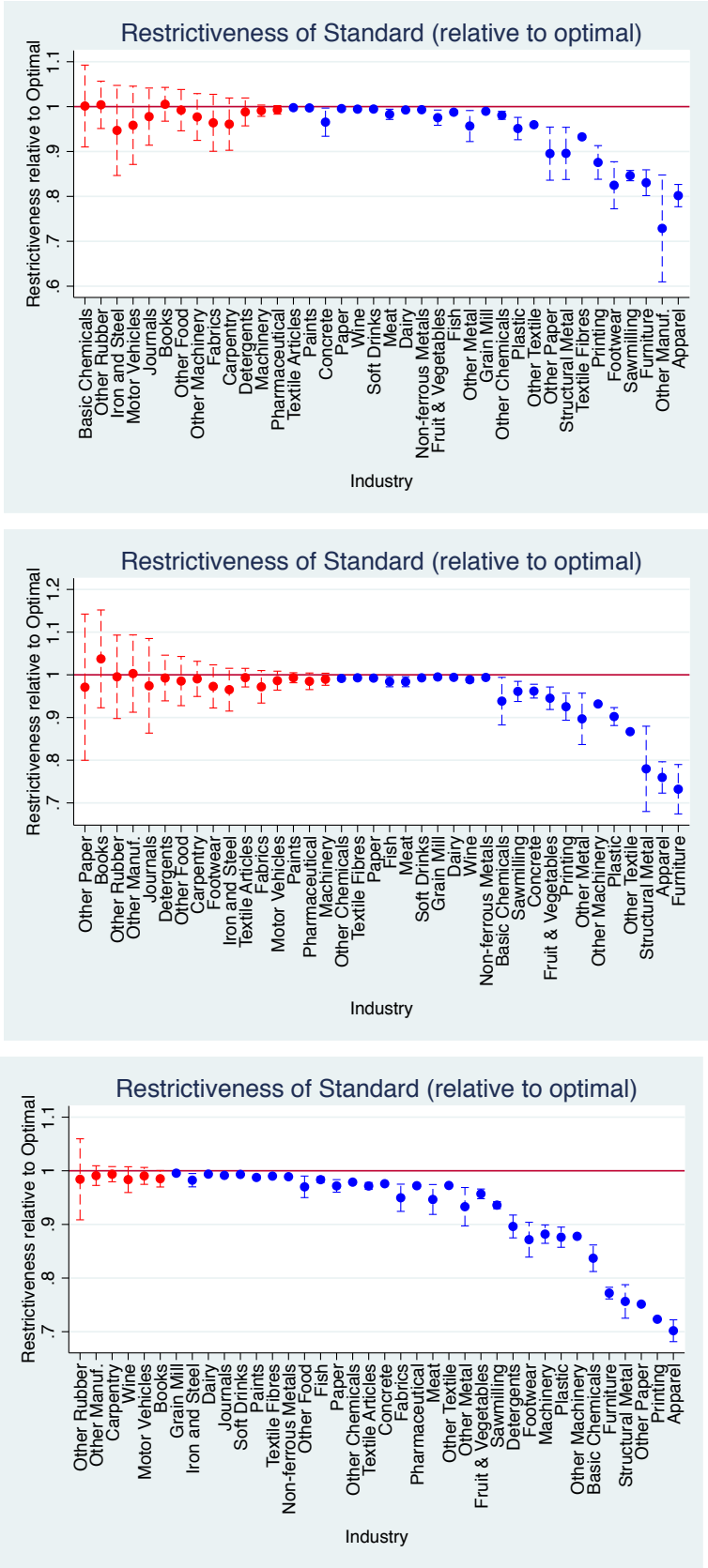






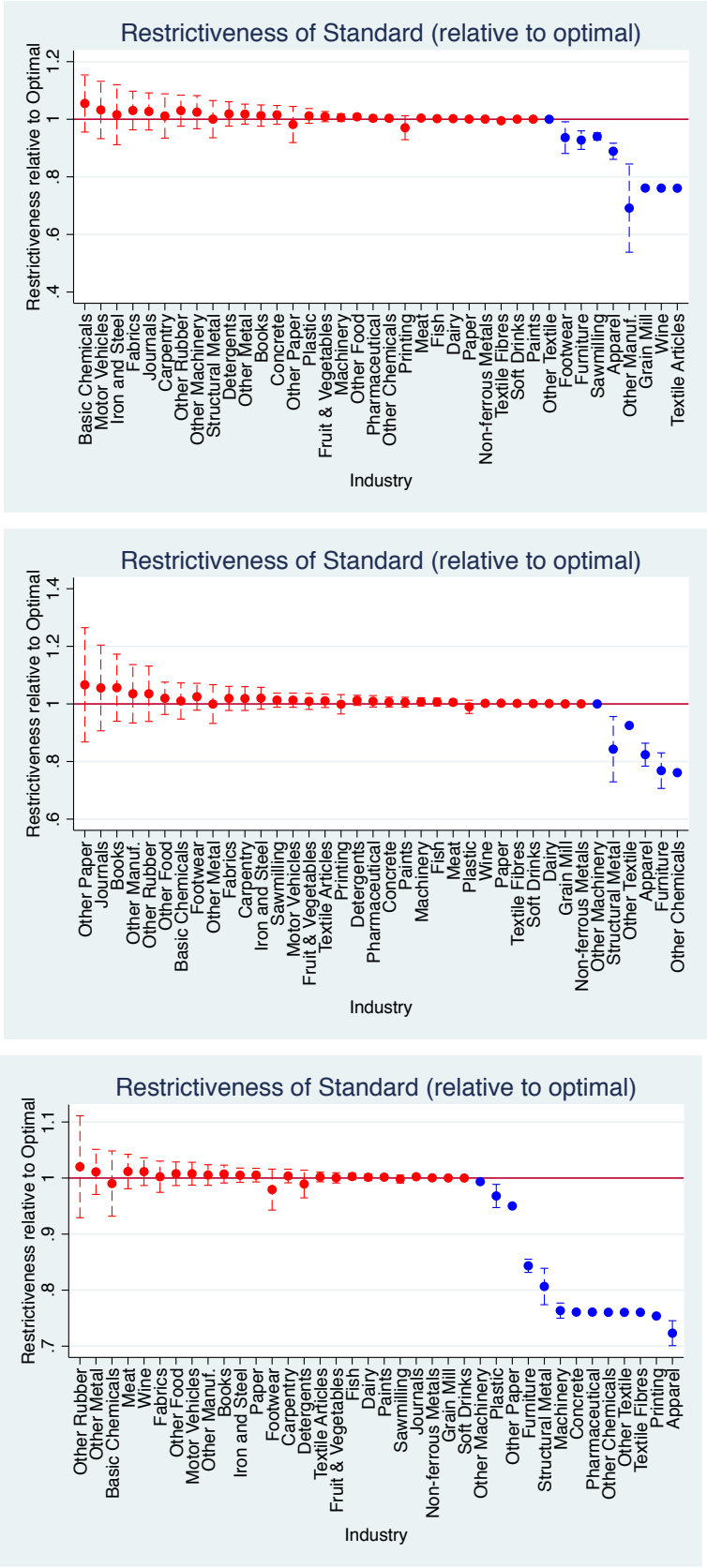
In red are the industries where we cannot reject the null hypothesis that the restrictiveness index is different from one.

Figure 17: Restrictiveness Index in DA with general γ Model: 1995, 2000, and 2005.



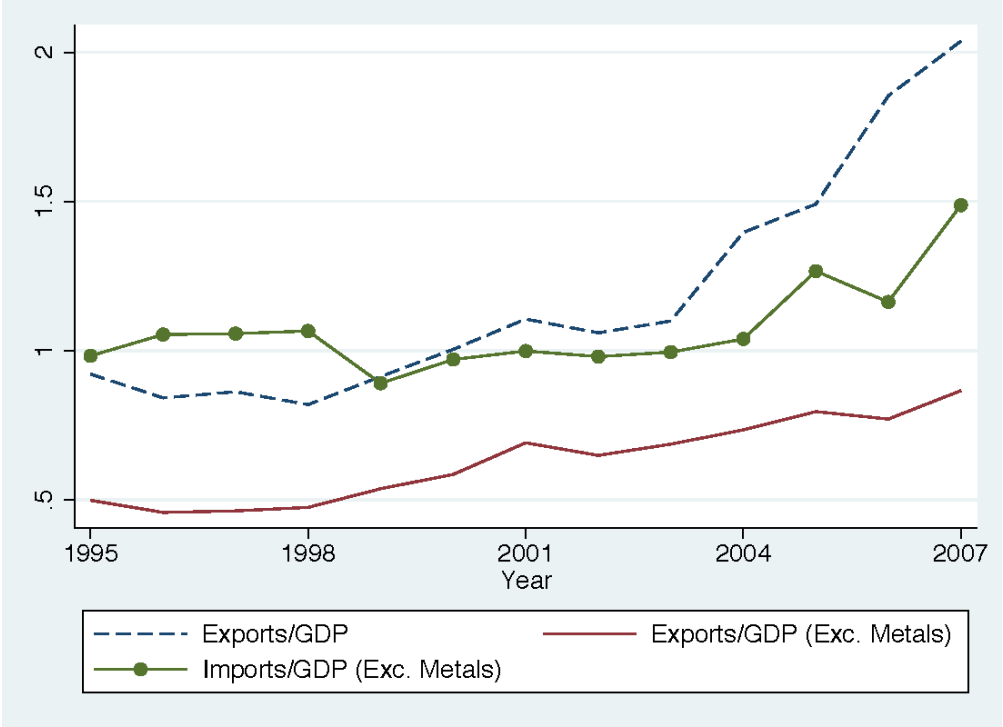
In red are the industries where we cannot reject the null hypothesis that the restrictiveness index is different from one.

Figure 18: Restrictiveness Index in IA with Fixed Costs: 1995, 2000, and 2005.



In red are the industries where we cannot reject the null hypothesis that the restrictiveness index is different from one.

Figure 19: Chilean Trade Flows, Tariffs, Terms of Trade



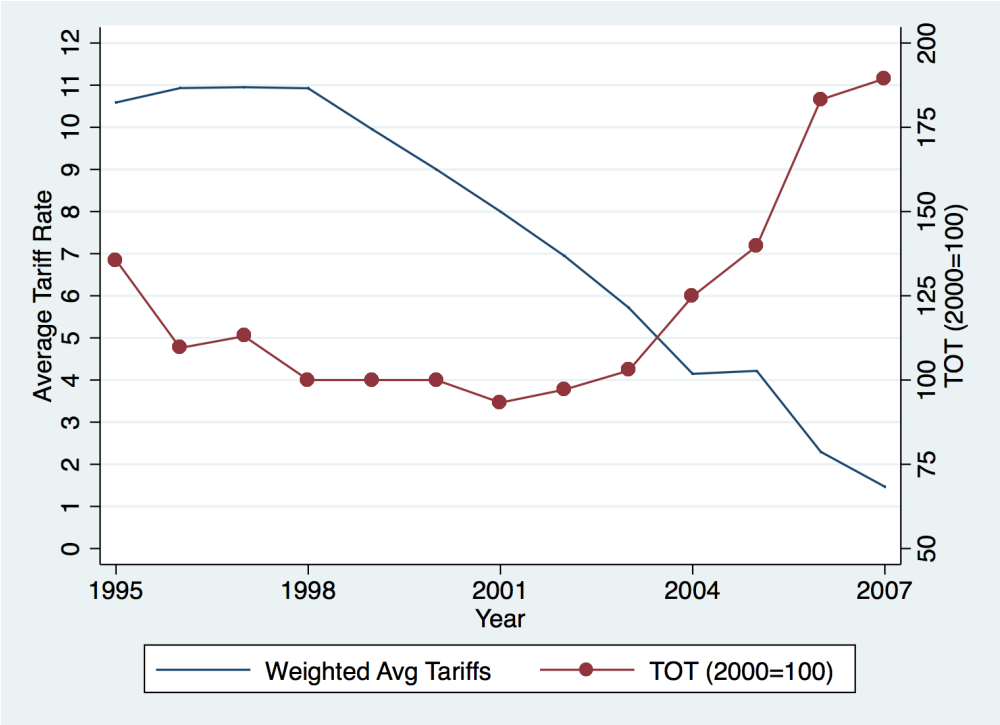
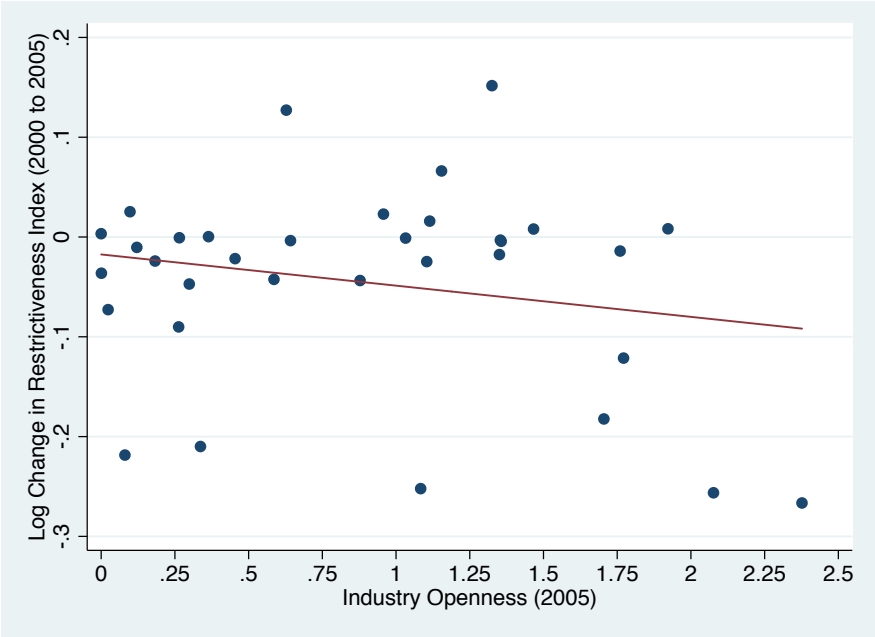


Figure 20: Industry Openness vs Log Difference in Industry Restrictiveness (2000 to 2005)



This figure plots Industry Openness versus the log change in the restrictiveness index between 2000 and 2005. Openness (x-axis) is defined as the sum of imports and exports over total sales. The y-axis is the log change in the restrictiveness index between 2000 and 2005. There are a total of 34 industries in the plot, as we drop industries with openness above a ratio of 4. The result is robust to allowing for industries with an even larger ratio, but we believe a cutoff is necessary because, for example, the "Other Manufacturing" industry has an openness ratio more than 60 times larger than the median industry. The slope of the best fit line is -.2.