Protectionism Unchained: Determinants and Consequences of Discretionary Trade Policy in Argentina

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January 2022
Preliminary and Very Much Incomplete
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Abstract

Many countries utilize discretionary trade policies—those that target particular firms at particular times—with such policies only increasing in their popularity given the recent backlash against globalization. However, since these types of policies are usually illegal under WTO rules, our ability to empirically study the forces that drive these episodes as well as their consequences has been severely hampered. To make progress, we study the prominent case of Argentina, a country that in the early 2010s implemented a set of discretionary trade policies that flagrantly violated WTO rules yet meticulously collected data on every government approval decision and trade flow. We first show that the government penalized larger import requests as well as imports into sectors with lower shares of capital importers, lower shares of exporters, lower average revenues and a higher fraction of domestically-owned firms. We then uncover a link between high-frequency variation in macroeconomic conditions and the level and structure of trade protection, with differences in protectionism across sectors narrowing with declines in foreign reserves as the government resorted to placing trade restrictions on initially-favored sectors. Second, we rely on this policy variation to identify the effects of import restrictions on import quantities and prices. As expected, import penalties reduced imports. But, surprisingly, the policy backfired through higher import prices at the port. A simple model of bargaining between foreign producers and importers explains such a result when foreign producers have most of the bargaining power—a scenario common to many developing-country imports. Estimates of the relative bargaining power across sectors provide support for the model which we then use to measure the aggregate impacts of these trade restrictions.

1 Introduction

Researchers have long theorized that governments may operate discretionary trade policies that favor or punish particular firms and sectors at different times. The recent backlash against globalization has only served to embolden governments in this regard (Baldwin and Evenett, 2020). A shift towards protectionist policies has been a tenet of political movements across many countries since at least the mid 1990s, and further accelerated since the 2008 financial crisis (Walter, 2021; Colantone et al., 2021). However, since many of these policies are illegal under WTO rules, understanding their determinants and consequences have been mostly a theoretical endeavor or has relied on anecdotal accounts.[1]

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[1]For example, China has been accused of trade practices that violate WTO rules such as providing illegal subsidies (Barwick et al., 2019), demanding local content restrictions, and imposing IP transfer requirements on foreign entities. While specific examples on these practices are well known, detailed data on these measures is not available.

In this paper, we use a rich dataset on trade policy restrictions imposed by Argentina in the early 2010's. This episode is somewhat unique in that the Argentinean government not only flagrantly violated WTO rules, but also kept a detailed record of these actions. From 2012-2015, the government instituted a system of non-automatic import licenses whereby every firm wanting to make an import shipment had to formally request to do so. Each request was then approved or denied at the government's discretion; as it turned out, these denials were frequent and widespread with around 30% of requests denied. Moreover, denial rates varied greatly over time, across firms, and across sectors. Thus, the discretionary trade policies instigated during this episode are not only well documented but also quantitatively important. The cornerstone of this paper is the electronic records for the universe of requests and government approval decisions between 2013 and 2015. We combine these records with transaction-level trade data and industry information to study the drivers and consequences of Argentina's discretionary trade policy.

We first identify a salient feature of the policy: larger import requests were more heavily penalized through smaller approval probabilities. These size penalties varied widely across narrow sectors (HS4) and over time, and were eliminated when the opposition party unexpectedly won the presidency by a narrow margin in 2015. We exploit this variation in penalties across sectors and over time, including high-frequency (half-yearly) swings in trade policy during 2013-2015, as well as the elimination of the policy altogether in 2015.

Our analysis of these restrictions demonstrates a novel determinant of trade policy. We uncover a link between high-frequency variation in macroeconomic conditions and both the level and structure of trade protection. During this period, Argentina experienced substantial exchange rate volatility and dwindling foreign reserves. We show that trade protection, measured by the likelihood of approvals and the request size penalty, became more stringent in periods of higher macro instability. Moreover, this time variation interacted with more traditional determinants of protectionism. In the cross section, sectors with higher shares of capital importers, higher shares of exporters, and higher average revenues were more likely to see import requests approved, while sectors with a higher fraction of domestically-owned firms were systematically provided with more protection while these policies where in place. These results are consistent with traditional industrial policy motivations as well as political-economy determinants of trade policy such as government preferences for redistribution to specific groups. However, the relatively unfettered ability to import enjoyed by the more protected sectors was eroded in periods of higher macro instability as the government turned to restricting these sectors as well in the quest to shore up dwindling foreign reserves.

In addition, we find substantial variation in the levels of protection afforded to firms within a given sector. A variance decomposition shows that the firm's identity is the primary determinant of approval probabilities rather than a firm's sector or the time period. In particular, firm dummies account for 30% of the variation in approval rates in a regression that also includes month dummies and product dummies. In very-preliminary analysis, these firm-level approval probabilities are not strongly correlated with firm size or ownership status, likely because anecdotal evidence suggests that a number of large foreign firms struck deals with the government to ease import restrictions in exchange for increased FDI. In work in progress, we further explore this firm-level heterogeneity and its consequences.

We then turn to the effects of this discretionary trade policy. Regardless of the specific policy motivation, a central view in trade theory posits that the aggregate economic effects of trade policy materialize through terms-of-trade effects (Bagwell and Staiger, 1999). Despite this central role, there is limited evidence to date on the price effects of protectionist trade policy in general, let alone discretionary policies or quantitative restrictions of the sort practiced by Argentina. Indeed, most of the evidence consists of trade-liberalization episodes rather than shifts to protectionism and focuses on WTO-legal tariff barriers rather than non-tariff barriers that are commonly used to achieve short-term protectionist goals that would otherwise require renegotiating a trade

agreement.[2] An obvious exception is the recent US-China trade war, with Amiti et al. (2019), Fajgelbaum et al. (2020) and Cavallo et al. (2019) all finding that US tariffs did not impact their import prices, shifting the burden of the tariffs to US consumers.

The variation in the structure of trade policy across narrow sectors (due to their capital importing, their exporting, etc.) and over time (due to macroeconomic conditions) that we discuss above constitutes a source of plausible exogenous variation in trade restrictions. Using transaction-level import data, we estimate the effects of these restrictions on the imported quantities and the prices paid to foreign suppliers. As expected, more stringent import restrictions substantially reduced imported quantities. However, we also identify a positive effect of these restrictions on the import prices paid by domestic importers to foreign companies. In other words, due to these import restrictions, Argentinean importers ended up paying more for less.

The finding of import price increases is inconsistent with a perfectly competitive trading environment in which, as in our case study, the importing country decides the allocation of the import licenses.[3] A competitive open economy would be either unable to affect world prices through import restrictions or, if it was, it would lower those prices by reducing the international demand for imports. However, this result is consistent with foreign exporters that hold market power. Therefore, we posit a standard model of import bargaining, where the terms-of-trade effects of trade policy are mediated by bargaining between importers and exporters.[4] We show that, in such a model, the import price increases with the type of penalties imposed by Argentina only if the foreign bargaining power is large enough. More generally, the extent of the price increases from these policies reflects the bargaining power of foreign exporters. Intuitively, the larger the foreign bargaining power, the more the model behaves like one where the equilibrium lies on the domestic import demand curve, so that import restrictions operate like negative supply shocks that raise the foreign price.[5]

In work in progress, we rely on these properties to estimate differences in foreign bargaining power across sectors and products from the responses of import prices to the request size penalties. We find that these bargaining power estimates correlate positively with plausible proxies of market power and the strength of outside options, for example the number of sellers and buyers in very narrow product segments. We then use the model to measure the aggregate impacts of these discretionary trade restrictions. Specifically, the panel evidence on the impacts of trade restrictions on quantities and prices is qualitatively consistent with aggregate movements, as average import quantities fell and import prices grew during periods of increased protectionism. Using the model, we aggregate the price and quantity effects of the policies to measure how much of the observed time trends are explained by the import restrictions. And in work in progress we perform counterfactuals assessing what the effects would have been had the bargaining power of Argentinian importers been more similar to the bargaining power of importers in a similarly-sized but more developed country.

The paper is organized as follows. Section 2 reviews the macroeconomic context and the main features of trade policy in Argentina leading up to and during 2012-2015. Section 3 contains a first pass at the data.

[2]See Fajgelbaum and Khandelwal (2021) for a review of this evidence. While tariffs are the quintessential trade policy instrument, their relevance for trade policy diminished throughout the second half of the 20th century through international negotiations and regional agreements. Due to the constraints imposed by these agreements, and the US-China trade war notwithstanding, many protectionist measures take the form of non-tariff barriers, which are harder to observe and regulate.

[3]Khandelwal et al. (2013) study the quantity and price effects of lifting quantitative restrictions imposed to Chinese textile exporters. In their case, exporters had to bid for an export quota in China, driving up the marginal costs. They find lower export prices with the export quota removal, which is consistent with lower export costs in a competitive environment. No such licensing on the exporter existed in Argentina, where the export rights were assigned discretionally to the importing firms.

[4]In a limiting case where exporters hold all the bargaining power, the property that these restrictions would raise prices is similar to the texbook result that, with a foreign monopoly moving along the domestic demand curve, a binding quota may raise import prices under particular shapes of that demand curve (Helpman and Krugman, 1989). Here, the rents resulting from the import licenses are held by domestic importers who are granted the right to import. As foreign exporters do not have rights on those rents, a situation of bilateral monopoly more accurately represents the institutional environment than a situation of foreign monopoly.

[5]Several frameworks have included a bargaining problem between importers and exporters includes; see McLaren (1997), Ornelas and Turner (2008), Antràs and Staiger (2012), Grossman and Helpman (2020), and Alviarez et al. (2021), among others.

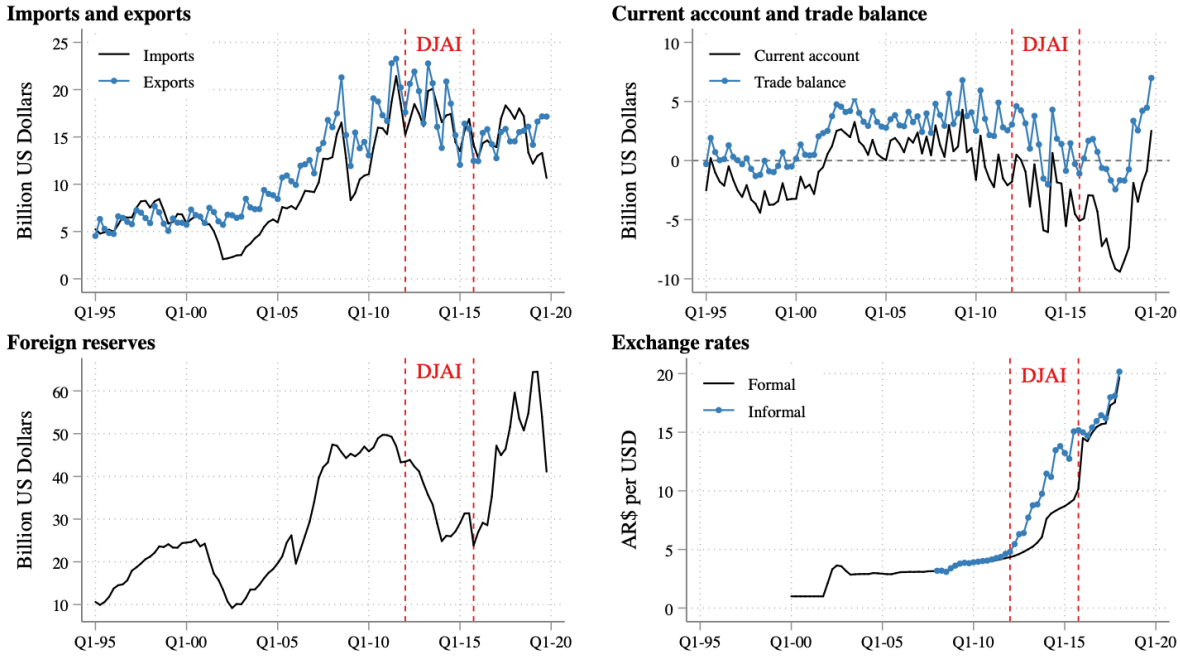
Section 4 presents the theoretical framework we use to discipline the approvals data and Section 5 empirically tests its main predictions. Section 7 concludes.

2 Institutional Background

In this section we provide a brief background to the discretionary trade policy—the Advanced Sworn Import Declaration system or DJAI by its initials in Spanish—that operated between 2012–2015.

After the crises of 2002, Argentina had almost a decade of strong economic growth. By 2009, in a context of a worsening current account and rising inflation (Figure 1, top right panel), the government began imposing restrictions to international trade. These policies came mostly in the form of non-automatic import licenses (NAILs) which, in contrast to automatic licenses, are not certain to be approved. WTO rules require that the basis for granting licenses be made public and their application be fair and non-discriminatory, conditions this policy did not satisfy. Between 2008 and 2011, the number of 8-digit products covered by NAILS grew from about 200 to 600 and the share of import value went from 5% to 17%—see Bernini and Garcia-Lembergman (2020). This licensing system was considered a Restrictive Trade-related Requirement by the European Union among others and led to disputes at the WTO (WTO 2014). By 2011, amid stagnating GDP and rapidly falling international reserves, the government implemented heavy restrictions on foreign currency purchases which led to the appearance of a black-market exchange rate (Figure 1, bottom right).

Figure 1: Imports, Exports, Reserves and Exchange Rate Leading to DJAI



Notes: The vertical lines indicate the first and last quarters when the DJAI system was in place. Top left: quarterly value imported and exported from FRED (Federal Reserve Bank of St. Louis). Top right: quarterly current acount balance and trade balance from INDEC (Argentinian National Institute of Statistics and Censuses). Bottom left: average quarterly stock of foreign currency reserves excluding gold from FRED. Bottom right: average quarterly exchange rate in the formal market (from BCRA, the Argentinian Central Bank) and in the informal market (from Ambito.com).

In February of 2012, the government launched a new system of trade interventions, the DJAI. The key difference with the previous policy was that the new regulations applied to all products imported into the

country. Firms had to request government authorization prior to making any purchase abroad, specifying detailed information such as value, quantity, origin and product HS code. Several agencies had the right to view these requests and place a given import request under review, the most prominent of which was the Secretariat of Trade. Requests under review could not be imported into the country as the Central Bank would not clear the payment and Customs would not release the goods.

When a request was under review, the importer had the option to appeal the decision with the relevant agency. In that case, the firm had to provide detailed information about its operation in the last three years as well as future projections. The information requested included imports and exports, domestic prices, and investments. It should be noted that the government did not officially publish any guidance or regulations detailing the grounds under which a request would successfully leave review status. However, importers were informally told that the Government's objectives were to improve the trade balance, to control domestic inflation and to foster import substitution and investments in Argentina. This review system came to a rapid end when the opposition party unexpectedly won the presidency by a narrow margin in 2015. Section A.2 of the Appendix contains additional details about the DJAI system. One of our goals in the next few sections is to shed more light on how this discretionary trade policy system operated in practice, including the firm-level and macro-level determinants of whether a request was placed under review.

3 Descriptive Evidence

In this section we document key properties of the transaction-level import restrictions implemented in Argentina. First, in Section 3.3 we show that the government frequently rejected import requests, a practice that all but disappeared when a new government was elected. We then show that approval rates were systematically related to the size of the request. Next, in Section 3.4, we establish a link between macroeconomic conditions and policy stringency, and identify sectoral characteristics associated with higher protection. These correlations form the basis for our later instrument for trade restrictions. Finally, in Section 3.5 we study the aggregate behavior of import quantities and prices over time.

3.1 Data Sources

We have records for all *import requests* filed from January 2013 to December 2017. The DJAI system was implemented informally in February 2012. Formal records started in 2013 and the policy was in place until late 2015, when a new government was elected. From January 2016 onwards, import requests within the DJAI system were still needed but were almost always approved, as we show. We obtained access to confidential data on the universe of these records. For each request we observe a firm tax identifier, the value and quantity requested, an 11-digit HS product code, the country of origin, and the measurement unit. Crucially, we also observe the quantity and value that was approved.[6]

In addition to the requests, we also observe actual trade transactions. Specifically, we observe the quantity and value imported and exported for the universe of transactions. These data are at the same level of aggregation as the requests and report the same firm and product information, hence we are able to verify the match between approved and realized transactions.[7]

[6]The product code is an 11-digit code that combines an 8-digit code from the NCM (Nomenclatura Común del Mercosur), and an extra 3-digit code specific to Argentina. The NCM is based on the Harmonized System (HS). For example, the 8-digit code 0901.11.10 contains coffee grains that are not roasted nor decaffeinated. The 11-digit codes split the beans further into arabica (0901.11.10.100) and robusta (0901.11.10.200).

[7]We further validate our trade data against the OECD and find that they match almost exactly (See Figure A.1).

We match the firm-level identifiers to D&B Hoovers data to obtain information about their industry (8-Digit SIC Code), sales, employment and ownership structure. We were able to match 68.5% of the firms in the sample to D&B, which amounts to 94.1% of the value imported in the sample period. We use the Ultimate Parent Company (UPC) to identify ownership for the 7.5% of our firms and 72.4% of import values where a UPC is reported and assume firms are domestically owned if they appear in D&B but have no UPC. This leads us to classify 5.95% of firms and 47.7% of import value as foreign-owned.

We obtain additional information on employment, compensation, and number of firms across 311 sectors (4-digit ISIC Rev. 3 codes) from the Observatorio de Empleo y Dinámica Empresarial (OEDE). These statistics are calculated using administrative social security data for the universe of formal private-sector employees in the country. From Datamyne and Comtrade we obtain series of values and quantities imported and exported at the 11-digit HS code level for additional years not included in our transaction-level data.

3.2 Approval Rates across Requests and Time

Table 1 presents summary statistics for approval rates across requests, and Figure 2 plots the time series of aggregate monthly values requested and approved. While the DJAI system was in place, about 30% of the requested value was denied. Shortly before the runoff election that decided the presidency on November 22, 2015, the incumbent government started approving all requests, and continued to do so until the change of government. On December 21, 2015, ten days after taking office, the new government ended the DJAI system, switching into a system of automatic licenses. That is, under the new government, importers still had to submit requests but those were almost always approved.

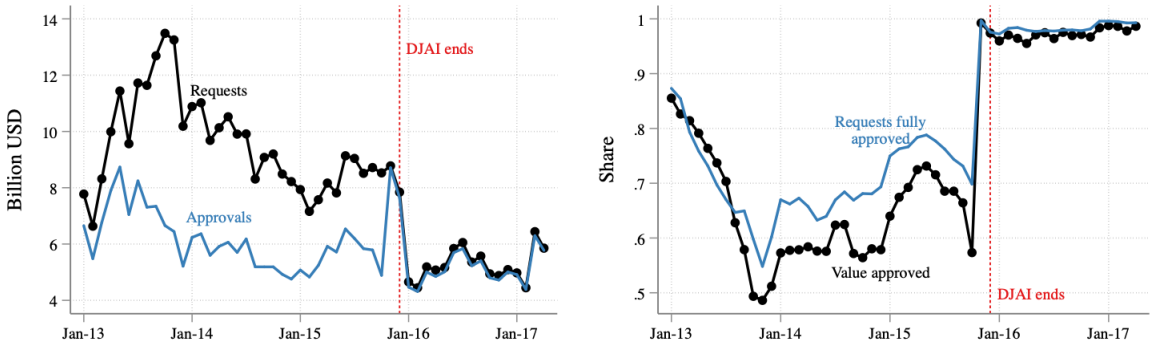
Figure 2 shows a sudden drop in the value and number of requests when the policy ended. This drop is consistent with importers inflating their requests, so a question arises as to whether the policy was binding or whether importers inflated the requests just enough to obtain their desired level of imports. It is also consistent with the end of the policy coinciding with an aggregate shock to desired imports. As preliminary evidence that neither of these were the case, Appendix A.4 shows that plausibly exogenous increases in approved quantities translated one-to-one to increases in imports. If the policy was non-binding, higher approved quantities would not translate into higher imports, as importers were not obliged to import the full amount of their request after it was approved. In addition, we show in Section 3.5 and in our main empirical analysis in Section 3.5 that differences in the stringency of the policy across products impacted quantities and prices as would be expected if the policy was binding.

Table 1: Descriptive Statistics of Requests and Approvals

	DJAI	Post DJAI
Requests per year	3,413,878	2,623,489
Mean value	$33,937	$26,277
Mean value approved	70.1%	98.5%
Total value approved	63.5%	89.5%
Requests fully approved	69.5%	86.4%

Notes: The DJAI period (first column) uses data from January 1st 2013 to October 1st 2015. The Post-DJAI period uses data from January 1st 2016 to December 31st 2017.

Figure 2: Total Value Requested and Approved

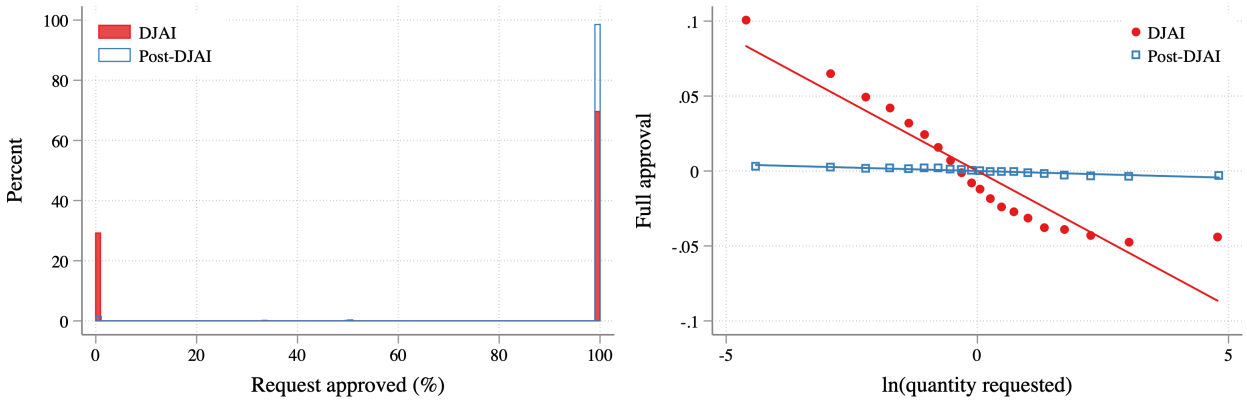


3.3 Import Restrictions by Request Size

We start by documenting two key features of the approval distribution which we will guide our analysis. First, most requests were either fully approved or fully denied. Figure 3 (left panel) shows the histogram of the share of value approved across requests. During the DJAI period (solid red), 69.5% of requests were approved fully, 29.2% were denied fully, and 1.3% were partially approved. In 2015-2017 (blue), when the policy was lifted, 98.4% were fully approved.

The second feature is that the likelihood of a full approval was systematically related to the size of the request. Figure 3 (right panel) contains a binned scatter plot of an indicator of full approval on the log quantity requested, after residualizing both variables on product dummies. In the DJAI period, the probability of a full approval was strongly and negatively related to the size of the request. When the policy was lifted, this relationship disappeared.

Figure 3: The Distribution of Approvals



Notes: The left panel shows a histogram of the percentage of quantity requested that was approved during the DJAI period and after the policy ended. The right panel shows a binned scatter plot and a linear fit of the probability of full approval on the log of the quantity requested, after residualizing both variables on product indicators. We define a product as a 11-digit category, origin, and unit triplet.

3.4 Import Restrictions, Sectoral Characteristics, and Macro Imbalances

We now show that trade policy restrictions varied across sectors and over time. These correlations suggest that the structure of trade policy was consistent with informal accounts from government officials regarding

the goal of the discretionary trade policy restrictions.

First, we show that the trade restrictions varied across products. In particular, we identify a number of sectoral characteristics that were systematically related to the level of protection. We consider the fraction of firms that were importers of capital goods, the fraction of exporters, the fraction of importers that were domestically owned, the log of average revenue and the log of average number of employees.[8] We define capital goods with the correspondence tables from 6-digit HS codes to broad economic categories (BEC) produced by the United Nations.[9] We work at the HS4 product level and measure these variables in 2011, prior to the enactment of the DJAI to alleviate endogeneity concerns (with the exception of revenues and employment which come from an October 2021 pull from D&B Hoovers and so should be treated with more caution). Column 1 of Table 2 regresses the approval rates by sector over the whole DJAI period on these characteristics. We find that sectors with a higher share of capital importers, a higher share of exporters, higher revenues and a lower fraction of domestically-owned firms were more likely to have their request approved. These finding are consistent with the government's rhetoric of encouraging domestic investment, promoting exports, and protecting domestic producers from import competition.[10] Column 2 regresses the sectoral approval rate in the first 6 months of the regime on the same characteristics with very similar results. The predicted values in both these columns will prove useful in constructing instruments to assess the impacts of Argentina's trade restrictions in Section 5.2.

Second, we show that the level and structure of protection correlated with macro imbalances. Starting in 2012, Argentina faced mounting pressure on its formal exchange rate. To defend the currency, the Central Bank intervened by selling international reserves and imposing exchange rate controls. As a result, foreign reserves fell and the black market exchange rate premium rose. During this period, we also observe that the government tightened controls on imports. Figure 4 shows a positive co-movement between international reserves and the fraction of requested value that was approved at a weekly frequency. This correlation suggests that trade policy restrictions were used as a tool to stabilize macroeconomic outcomes alongside reserve interventions and exchange rate controls.

As we document in Section 5.2.2, these twin determinants of Argentina's trade policy in this period—firm and product characteristics as well as macro imbalances—interact with each other during the DJIA period. In particular, the sectors less heavily targeted at the start of the DJIA regime (i.e. those with characteristics associated with high approval rates) experienced larger reductions in approval rates as international reserves fell. In effect, the government appears to have an ordering of which sectors to target and was forced to go further down the list and also restrict imports in more favored sectors when foreign reserves were particularly scarce.

[8] All fractions and averages use 2011 import values as weights. Appendix Figure A.6 shows that the domestic ownership share is positively correlated with employment, the number of firms, and the wage bill across sectors.

[9] The correspondence table is available at `https://unstats.un.org/unsd/trade/classifications/correspondence-tables.asp`. We consider as capital goods the 6-digit HS codes classified as capital goods, and as non-capital goods every other HS code (including the unclassified ones). With this classification, 72.0% of firms imported a capital good at least once and 20.7% of total import values were capital goods.

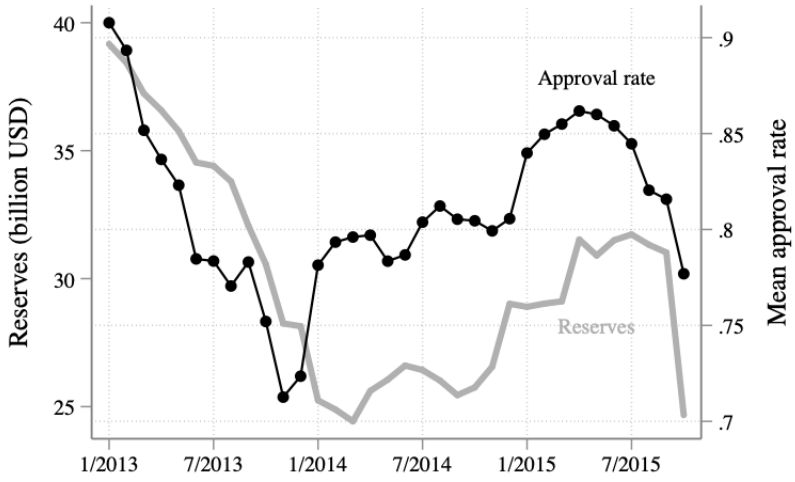
[10] The WTO panel documented public statements and speeches where high-ranking officials announced a policy of "managed trade" with the objectives of substituting imports for domestically-produced goods and reducing or eliminating trade deficits. They also found evidence that the DJAI system was a tool of Argentina's "managed trade" policy (WTO, 2014).

Table 2: Approvals and sectoral characteristics under the DJAI system

	AR	$AR^{\text{H1-13}}$
Fraction of Capital Importers	0.057*	0.100***
	(0.022)	(0.027)
Fraction of Exporters	0.137***	0.158***
	(0.025)	(0.031)
Fraction Domestically Owned	-0.076***	-0.116***
	(0.019)	(0.025)
log(Mean Revenue)	0.017*	0.022*
	(0.007)	(0.010)
log(Mean Employees)	-0.012	-0.019
	(0.009)	(0.010)
Constant	0.479***	0.351*
	(0.098)	(0.137)
N	1,102	1,111
R^2	0.169	0.191
F-stat	33.2	42.2

Notes: Table shows regressions on HS-4 product-level approval rates on sectoral characteristics. The approval rate AR is the HS4-level mean of the ratio of quantity approved to requested during the whole DJAI regime, January 2012 to October 2015 (column 1), or during the first 6 months of the regime, January–June 2013 (column 2). Robust (HC3) standard errors shown in parentheses. Asterisks indicate 10% (*), 5% (**), and 1% (***) significance.

Figure 4: Approvals and foreign currency reserves under the DJAI system



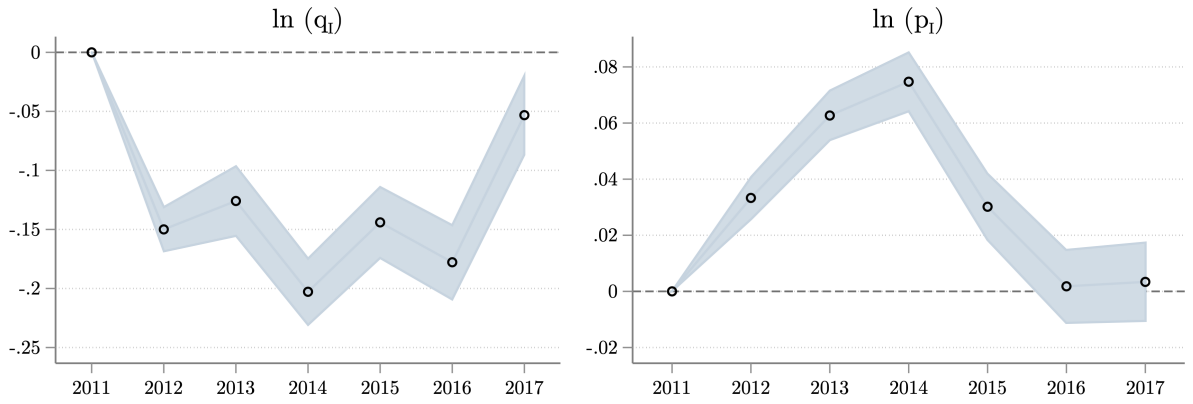
Notes: Figure plots both approval rates and foreign currency reserves over time. Reserves are monthly averages of the level of foreign currency reserves (solid gray line). The approval rate, plotted as a dotted line, is the total monthly approved import value divided by the total monthly requested import value.

3.5 Import Prices and Quantities

We now turn to the two key outcomes of interest in the paper, import quantities and prices. As a first-pass at the effects of the DJAI system, we run regressions of log import quantities and prices on year dummies and firm-product fixed effects. Figure 5 plots the values of the recovered year dummies. We find that the import restrictions were associated with a large decrease in import quantities and an increase in import prices, a

relationship that holds even if we look within narrow firm-product combinations. Within a firm and product, quantities fall on average by 0.2 log points (18.13%) from 2011 to 2014 and prices rise by 0.075 log points (7.79%). After the DJAI policy ended in late 2015, imported quantities increase but still remain 0.05 log points (4.88%) below the level in 2011, while prices return to their 2011 levels.

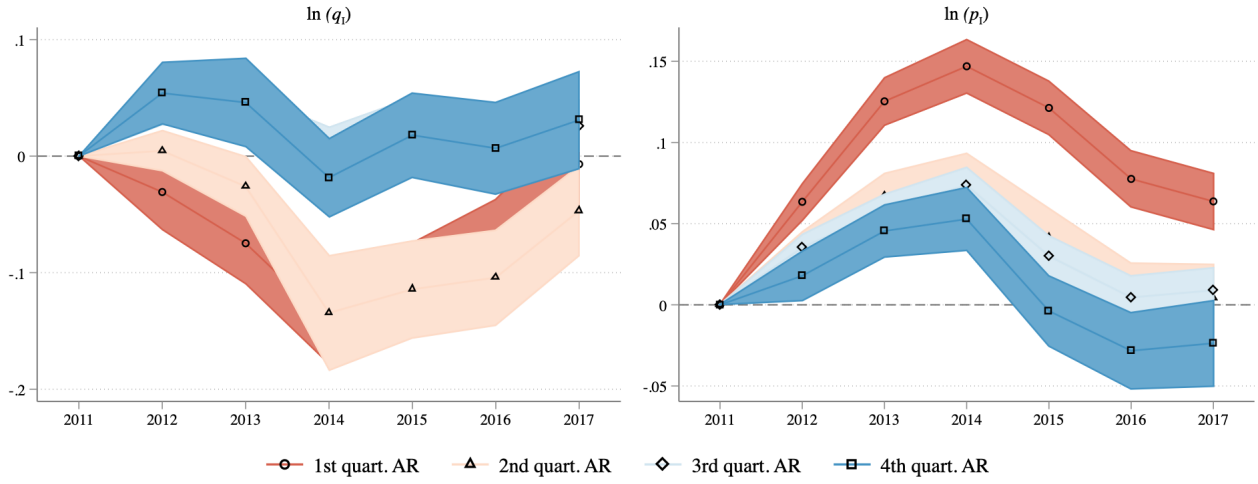
Figure 5: Price and quantity changes within firm-product



Notes: Plot of the μ_t fixed effects from estimating $y_{fit} = \mu_t + \mu_{fi} + error_{fit}$ where i and f are product and firm identifiers. The variable y is log imported quantities (q_I) in the left panel and log import prices (p_I) in the right panel. The standard errors are clustered by firm. The data is aggregated to a yearly frequency.

Are these patterns related to the DJAI system? As a first pass, we explore the behavior of quantities and prices of sectors with varying degrees of protection. We group the 4-digit products into 4 categories depending on their mean approval rate (value approved over value requested) during the DJAI period. We then regress log import quantity and price on firm-product fixed effects and year dummies for each product group, and plot the time dummies (Figure 6). The figure reveals that sectors where the policy was more stringent feature lower import quantities and higher prices during the DJAI period.

Figure 6: Within firm-product price and quantity changes by average approvals during DJAI



Notes: Plot of the μ_t fixed effects from estimating $y_{fit} = \mu_t + \mu_{fi} + error_{fit}$ where t is a year, i a product, and f is a firm. Regression run separately by groups of HS4 sectors determined by their average approval rate during the DJAI system. The membership to a group is fixed in time. The LHS variable is log imported quantities (q_I) in the left panel and log import prices (p_I) in the right panel. The standard errors are clustered by firm.

To provide a more complete analysis, the next sections present a theoretical framework linking firm-level trade policy to import quantities and prices which will allow us to formalize an instrumental variable strategy that leverages the variation described in Section 3.4 and provide supporting evidence for the mechanisms at work.

4 Theory

This section proposes a model of a transaction between a domestic and a foreign firm subject to regulations akin to those of Argentina in 2012-2015.

4.1 Environment

The domestic firm can buy a good from the foreign firm and use it as the single input in production with a linear technology. Letting q be the quantity imported, the domestic firm faces demand:

$$p(q) = aq^{-1/\sigma},$$

where $a > 0$ is a demand shifter and $\sigma > 1$ is the elasticity of demand. The foreign exporter has a per-unit production cost given by

$$c(q) = zq^{\frac{1}{\eta}}$$

where z is a cost shifter and $\eta \geq 0$ is the elasticity of supply.

In order to complete an international transaction, the firms must obtain government authorization. The timing is as follows. The importer-exporter pair first submit a quantity request. The Government then decides the fraction of the request that is approved. The firms decide an import quantity and engage in Nash bargaining over the price.

Motivated by the features of the Argentine policy documented in Section 3 above, we make the following two assumptions about the approval process. First, the request is either fully approved or fully rejected. Second, the likelihood of full approval is decreasing in the size of the request. In particular, letting q_R denote the quantity requested, the probability of full approval is given by:

$$\alpha(q_R) = \phi_0 q_R^{-\phi},$$

where $\phi_0, \phi > 0$. Here, ϕ controls the elasticity of the full approval probability to the size of the request. We refer to it as the request size penalty parameter.

4.2 Optimal Price and Quantity

We solve the model backwards. Given an agreed import quantity q, the per-unit import price is determined by:

$$p^*(q) = \arg\max_T \left[(p(q) - T)\, q\right]^{\beta} \left[(T - c(q))\, q\right]^{1-\beta},$$

where $\beta \in (0, 1)$ is the Nash bargaining weight on the importer. It is straightforward to show that

$$p^*(q) = (1-\beta)\, aq^{-\frac{1}{\sigma}} + \beta z q^{\frac{1}{\eta}}. \tag{1}$$

Intuitively, the bargained transfer is a weighted average of the importer's revenue per-unit and the exporter's average cost, with weights given by the bargaining parameter. A high bargaining weight of the exporter (low

β) pushes the transfer closer to the revenue per unit. It follows that the bargained price can be decreasing or increasing in the quantity imported, depending on which of these components dominates.

Consider now the choice of the quantity to request given the bargained transfer function. Regardless of which party decides, the optimal quantity will maximize total expected surplus:

$$q_R = \arg\max_q \ \alpha(q)\left(p(q) - c(q)\right)q \tag{2}$$

Claim 1. When $1 - \frac{1}{\sigma} > \phi$, the profit maximization problem in the first stage has a unique solution given by

$$q_R = \left(\frac{a}{z}\frac{1 - \frac{1}{\sigma} - \phi}{1 + \frac{1}{\eta} - \phi}\right)^{\frac{1}{\frac{1}{\eta} + \frac{1}{\sigma}}}. \tag{3}$$

Proof. See Appendix. □

Expressions (3) and (1) fully characterize the optimal quantity requested and import price as functions of parameters.

4.3 Predictions

We now derive three model predictions that we will later test empirically.

Claim 2. An increase in the request size penalty parameter leads to a reduction in the quantity requested:

$$\frac{\partial}{\partial \phi} q_R < 0.$$

Intuitively, by reducing the probability of approval for larger requests, a higher ϕ reduces expected surplus more strongly for higher values of q. In contrast, ϕ_0 affects expected surplus equally across all values of q and therefore has no effect on the optimal quantity requested. We now turn to the effect of the request size penalty parameter on the optimal import price.

Claim 3. The import price is increasing in the request size penalty parameter whenever the foreign supplier has enough bargaining power:

$$\frac{\partial}{\partial \phi} p^* > 0 \text{ and } \frac{\partial}{\partial q} p^*(q) < 0 \iff \frac{\beta}{1-\beta} \leq \frac{\eta + 1 - \eta\phi}{\sigma - 1 - \sigma\phi} \tag{4}$$

Thus, a tighther penalty parameter can increase or decrease the import price depending on parameters. The parameter condition in (4) is intuitive. The left hand side captures the relative bargaining power of the importer. The right hand side captures the "technological" relative power of the foreign supplier. An inelastic domestic demand (low σ) or an elastic foreign supply (high η) gives more power to the foreign supplier. The import price is increasing in the penalty parameter whenever the seller is more powerful, either in terms of barganing or technologically.

We next explore whether the effect of the penalty parameter on import prices and quantities depends on the bargaining power.

Claim 4. Regardless of its sign, the change in import price following an increase in the penalty parameter is higher when the foreign supplier bargaining power is high. In contrast, the effect of the penalty parameter on

the quantity requested is independent of the barganing weights:

$$\frac{\partial^2}{\partial\beta\partial\phi}p^* < 0 \text{ and } \frac{\partial^2}{\partial\beta\partial\phi}q_R = 0.$$

5 Empirics

We now test the model predictions derived above by exploiting variation across products and time. We start by estimating the request size penalty parameter, ϕ, by detailed product category. We then study the effect of changes in the estimated $\phi's$ on import prices and quantities, and how these effects depend on bargaining power. We focus on two types of variation in the size penalty parameter: between policy regimes (pre, during, and post DJAI) and within the DJAI period, with the latter providing more convincing causal estimates for reasons we discuss below.

5.1 Estimating the Size Penalty

We estimate an approximation to the size penalty in a panel regression that leverages the full granularity of our data. In our preferred specification, we estimate the following model:

$$\mathbb{1}\{q^A_{fit} = q^R_{fit}\} = \phi_{0,h} - \phi_h \ln(q^R_{fit}) + \mu_f + \gamma_t + e_{fit}, \tag{5}$$

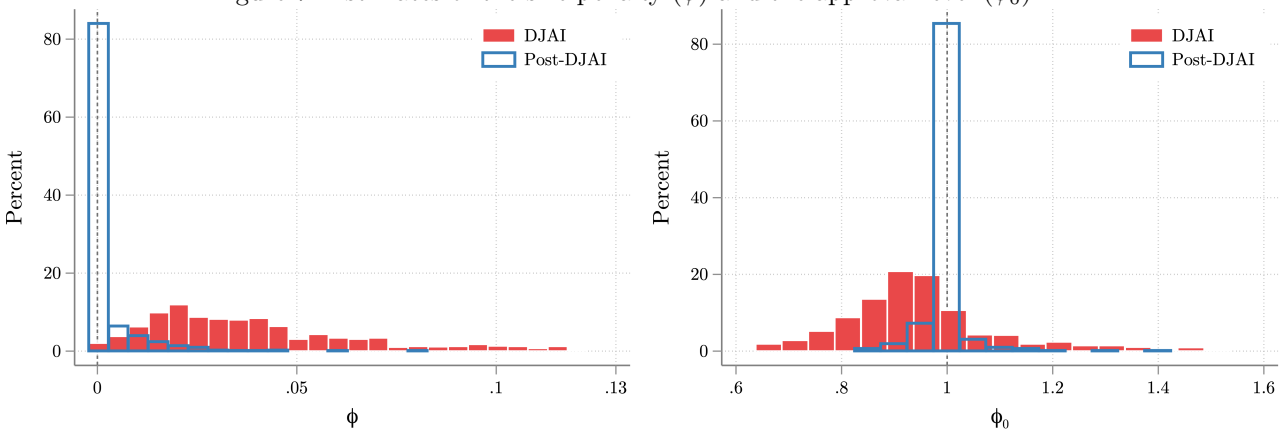
where q^R is the quantity requested, q^A is the quantity approved, and $\mathbb{1}\{q^A_{fit} = q^R_{fit}\}$ is a dummy variable that takes the value 1 if the full request was approved. In terms of subscripts, f identifies firms, i identifies products (11-digit HS - origin - unit of measurement triplets), h identifies a 4-digit HS code, and μ_f and γ_t are firm and month-of-sample fixed effects. The parameters of interest, ϕ_h, are the semi-elasticities of the probability of full approval to the size of the request. The fixed effects imply that, for each 4-digit sector h, we are using variation in approval sizes within the same firm and the same month of the sample. We estimate (5) separately for the two policy regimes where requests data exist: January 2013 to October 2015 and January 2016 to December 2017. That is, we estimate two ϕs and two ϕ_0s for each 4-digit HS category. Section 5.2.2 performs a similar exercise to extract variation in ϕs within the DJIA period.

Figure 7 plots the distributions of the parameter estimates. Our estimates imply that during the DJAI period there was a penalty for the size of the request that disappeared when the policy ended (left panel). Moreover, there is significant variation between 4-digit product categories in the magnitude of that penalty. During the DJAI period, we estimate a positive average size penalty of 0.043, with bottom and top quartiles of 0.019 and 0.056, respectively. In contrast, after the DJAI ended the distribution of the size penalties virtually collapses to having all the mass at zero. The estimates of the approval level parameter ϕ_0, depicted on the right panel, also feature dispersion across products in the DJAI period. The average is 0.97 and the bottom and top quartiles are 0.86 and 1.00, respectively. After the policy ended, the distribution collapses to having virtually all the mass at one.

The patterns in Figure 7 are robust in a number of ways. First, most of the estimates are significantly different from zero at conventional levels. Second, the differences between the DJAI period and the post-DJAI period hold if instead we estimate the parameters separately for every year in the sample or for every quarter in the sample. Third, the pattern is also similar if we estimate Equation 5 but with a fully-saturated set of fixed effects. The details are in the Appendix.[11]

[11] Figure A.3 plots the estimates together with confidence intervals and Figure A.2 plots the distribution of the estimates by year. Figure A.4 shows the distribution of the estimates by quarter-of-sample. In Table A.1 we show statistics of the distributions of the estimated parameters if we keep only significant estimates and if we use a fully-saturated version of Equation 5.

Figure 7: Estimates of the size penalty (ϕ) and the approval level (ϕ_0)



Notes: Histograms of the estimates of ϕ (left panel) and ϕ_0 (right panel) at the level of the 4-digit HS code during the DJAI policy (2013-2015) and after the policy (2016-2017). Top and bottom 2% excluded for readability.

5.2 Size Penalty, Import Prices and Quantities

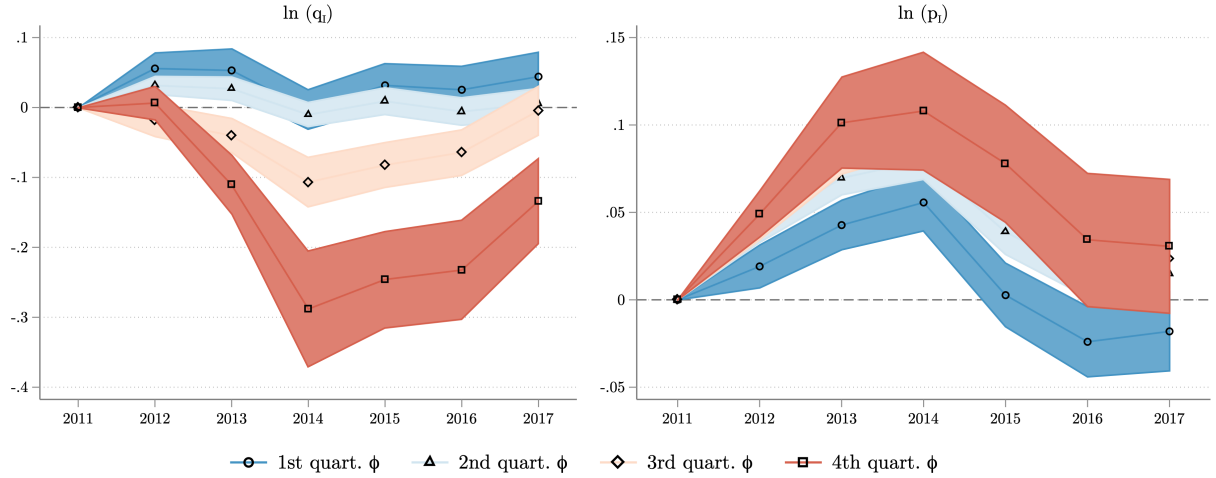
Armed with the policy parameter estimates, we now assess the model's predictions and use the estimated quantity and price responses to learn about the underlying parameters. We focus on the relationships between the request size penalty, import quantities and prices. According to the model, an increase in ϕ should unambiguously lead to a reduction in import quantity while the effect on the price depends on parameter values (Claims 2 and 3). We first exploit changes between policy regimes and repeat the analysis using quarterly variation within the DJAI period in Section 5.2.2.

5.2.1 Policy Regime Changes

We start by exploring graphically the behavior of import prices and quantities for products with different size penalties. To do so, we group the 4-digit products into quartiles of the estimated ϕ during the DJAI period. We then regress log import quantity and price on firm-product fixed effects and year dummies for each of the four quartiles, and plot these recovered year dummies by quartile (Figure 8). We find that sectors where large requests were more-heavily penalized displayed lower import quantities and higher import prices during the DJAI period.[12] This is consistent with the model when the foreign supplier's bargaining power is high (Claim 3).

[12]Table A.2 in the Appendix illustrates the diversity in the products included in the different categories. In the appendix, we also show the results of a specification where we run a single regression comparing the sectors above and below the median ϕ, and include HS category-by-time fixed effects (Figure A.7). The results are consistent with Figure 8.

Figure 8: Within firm-product price and quantity changes by the magnitude of size penalty



Notes: Plot of the year fixed effects μ_t from estimating $y_{fit} = \mu_t + \mu_{fi} + error_{fit}$ where t is a year, i a product, and f is a firm. Regression run separately by quartiles of HS4 sectors determined by the estimated value of ϕ. The membership to a group is fixed in time. The LHS variable is log imported quantities (q_I) in the left panel and log import prices (p_I) in the right panel. Standard errors clustered by firm.

To explore this hypothesis more directly, we turn to regressions that exploit the full variation across disaggregated product categories. Specifically, we regress the changes in prices for each firm-product on the change in quantities, instrumenting the latter with changes in the policy parameter $\Delta\widehat{\phi}_{ht}$. According to the theory, the relationship between prices and quantities induced by shocks to the size penalty is informative about the relative bargaining power of foreign suppliers. We will address the potential endogeneity of the policy shocks below but for now we treat them as valid instruments for changes in quantities. Specifically, we run the following specification:

$$\Delta \ln(p^I_{fit}) = \rho \Delta \ln(q^I_{fit}) + \mu_{fi} + \gamma_t + u_{fit}, \tag{6}$$

where p^I and q^I are import prices and quantities, respectively, and $t \in \{pre, during, post\}$ denotes whether we consider the pre-DJAI period, the DJAI period, or the post-DJAI period. As before, f identifies firms, i identifies products (11-digit HS - origin - unit triplets), h identifies a 4-digit HS code, and μ_{fi} and γ_t are firm-product and month fixed effects. The quantities q^I_{fit} and prices p^I_{fit} are averages across the transactions of a firm and product within a policy regime. We use changes in the size penalty $\Delta\widehat{\phi}_{ht}$ as an instrument for changes in imported quantities $\Delta \ln q^I_{fit}$ and assume that the size penalty was zero before the pre period, i.e., we set $\widehat{\phi}_{h,Pre} = 0$.

Table 3 reports the results of this regression. We find that the changes in the imported quantity are negatively associated with the changes in the size penalty parameter (the first stage of this IV regression, column 1), as predicted by the model. We also find that changes in quantity induced by changes in the penalty parameter are negatively associated to changes in price, i.e., $\hat{\rho} < 0$ (the second stage, column 4). In the theory, this is consistent with relatively high bargaining power of foreign suppliers (Claim 3). Reassuringly, an increase in ϕ is associated with an increase in import prices (the reduced form, column 3).

Instrumenting for Policy Changes So far, we have treated the product-level changes in size penalty as exogenous. A natural concern, however, is that policymakers chose the penalty ϕ based on the sector's import value. To partially address this endogeneity concern, we rely on the relationship between trade protection

Table 3: The effects on import prices and quantities across policy regimes

	(1) 1^{st} stage $\Delta\ln(q_I)$	(2) OLS $\Delta\ln(p_I)$	(3) Reduced form $\Delta\ln(p_I)$	(4) 2^{nd} stage $\Delta\ln(p_I)$
$\Delta\phi$	-2.425***		0.657***	
	(0.38)		(0.15)	
$\Delta\ln(q_I)$		-0.325***		-0.271***
		(0.01)		(0.06)
Firm-product FE	Yes	Yes	Yes	Yes
Period FE	Yes	Yes	Yes	Yes
Observations	385,984	385,984	385,984	385,984
Importers	10,927	10,927	10,927	10,927
F-stat	39.7			

Notes: Table reports regressions of price changes on quantity changes with quantity changes instrumented by changes in the estimated size penalty between policy periods. Standard errors clustered by firm shown in parentheses. Asterisks indicate 10% (*), 5% (**), and 1% (***) significance.

and the (predominantly pre-DJIA) sectoral characteristics documented in column 1 of Table 2 in Section 3.4. Recall that we found that protection was stronger in sectors with a lower share of capital importers, a lower share of exporters, lower revenues or a higher share of domestic ownership prior to the policy.

Motivated by these findings, we rely on sectoral characteristics to predict the changes in ϕ in entering and exiting the DJAI regime. The assumption here is that these characteristics were chosen for historic reasons (e.g. classic industrial policy ideas) and that the industries with these characteristics did not have differential trends around the time of the DJIA. If this assumption holds true, we can obtain the predicted approval ratios from column 1 of Table 2 in Section 3.4 and then use the interaction between this predicted approval ratio and a dummy for the DJAI period as an instrument for the change in the size penalty.[13] Specifically, we predict the changes in ϕ by interacting the change in an indicator of whether the DJAI regime is in place with the approval ratio predicted by the pre-policy characteristics regression:

$$\Delta\widehat{\phi}_{ht} = \kappa\Delta\mathbb{I}\left(DJAI\right) \times \widehat{ApprovalRatio}_h + \mu_{fi} + \gamma_t + v_{fit}, \tag{8}$$

where h is a 4-digit HS code and $\widehat{ApprovalRatio}_h$ is the predicted approval ratio obtained from (7) in footnote 13.

Armed with these predicted changes in $\phi's$, we repeat the analysis of Table 3. Table 4 contains the results of this three-stage approach. Of course, this is equivalent to a two-stage approach with $\Delta\mathbb{I}\left(DJAI\right) \times \widehat{ApprovalRatio}_h$ serving as an instrument for $\Delta\ln(q^I_{fit})$ but we believe the three stage approach makes the logic more transparent. First, we note that the DJAI indicator crossed with the predicted approval ratio instrument has predictive power for changes in the size penalty ('zero' stage, column 1): the rise of the DJAI system mapped into increases in the penalty parameter that were particularly large for sectors with low predicted approval rate. Thus, for example, sectors with low prevalence of capital imports in 2011 saw a larger increase in the penalty parameter ϕ under the DJAI regime. Using the predicted changes in ϕ from this zeroth stage, we report the first stage, reduced form, and second stage coefficients that are qualitatively similar to

[13] Column 1 of Table 2 in Section 3.4 regresses the DJAI approval ratio on the aforementioned characteristics measured prior to the DJIA:

$$ApprovalRatio_h = \delta X_{h,2011} + \epsilon_h, \tag{7}$$

where h is a 4-digit HS code, $ApprovalRatio_h$ is the ratio of value approved to value requested during the DJAI period, and $X_{h,2011}$ is a vector of pre-policy sectoral characteristics that includes the fraction of capital importers, the fraction of exporters, the fraction of domestically owned firms, log revenue and log employees.

those of Table 3. That said, this is ultimately a three-period panel regression and the identification assumption that sectors with these characteristics were not on different trends is not entirely credible, particularly since we select the weights on the different characteristics based on the DJAI approval rates. Thus we now turn to a more credible analysis that exploits variation within the DJAI period.

Table 4: The effects on import prices and quantities instrumenting for size penalties

	(1)	(2)	(3)	(4)	(5)	(6)
			1^{st} stage	OLS	Reduced form	2^{nd} stage
	$\Delta\phi$	$\Delta\ln(q_I)$	$\Delta\ln(q_I)$	$\Delta\ln(p_I)$	$\Delta\ln(p_I)$	$\Delta\ln(p_I)$
$\Delta\mathbb{1}\{DJAI\} \times \widehat{AR}$	-0.207***		0.613***		-0.111*	
	(0.012)		(0.095)		(0.054)	
Predicted $\Delta\phi$		-3.013***				
		(0.471)				
$\Delta\ln(q_I)$				-0.325***		-0.181*
				(0.015)		(0.078)
Firm-product FE	Yes	Yes	Yes	Yes	Yes	Yes
Period FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	385,984	385,984	388,872	388,872	388,872	388,872
F-stat	303.2	41.0	41.3			

Notes: Table reports regressions of price changes on quantity changes, with quantity changes ultimately instrumented by sector characteristics interacted with a dummy for the policy period. The standard errors are one-way clustered by HS4-period and shown in parentheses. Asterisks indicate 10% (*), 5% (**), and 1% (***) significance. The predicted approval rate, $\widehat{AR}$, is the fitted value from Table 2.

5.2.2 Variation within the DJAI period

So far we have relied on time variation between policy regimes. We now turn to higher frequency (6-month) variation within the DJAI policy period. As above, we can regress price changes on quantity changes but now just exploiting changes in price and quantities within the policy period when the DJAI was in place. And again we can use a three stage approach instrumenting the change in quantity with the change in the size penalty over time, with the change in size penalty itself a function of policy determinants.

To implement such an approach, we first re-estimate the size penalties, the 4-digit HS specific ϕs, separately for each half-year period. Figure A.5 in the Appendix plots the distribution of these estimates.

We then consider an instrument for the changes in the size penalties within the DJAI period that relies on the idea first previewed in Section 3.4 that the government increased trade protection when international reserves fell, and disproportionately so for sectors that it had not initially targeted. In particular, we predict half-yearly changes in $\phi's$ by interacting the change in international reserves with the predicted approval ratio based on (predominantly pre-DJIA) sectoral characteristics. To mitigate endogeneity concerns related to the government choosing characteristics based on sectors that were struggling at the time, we utilize the predicted approval rate based on the first sixth months of approvals, $\widehat{ApprovalRatio}_h^{\text{H1-13}}$, and look at half-yearly changes in subsequent periods (i.e., these predicted approval rates are formed from the coefficients reported in column 1 of Table 2 in Section 3.4.) Thus, the zeroth stage to predict the changes in the penalty parameter is now:

$$\Delta\widehat{\phi}_{ht} = \kappa\Delta\ln(Reserves_t) \times \widehat{ApprovalRatio}_h^{\text{H1-13}} + \mu_{fi} + \gamma_t + v_{fit}, \tag{9}$$

where h is a 4-digit HS code, t is a half-year period during DJAI (excluding the first half of 2013), and $\widehat{ApprovalRatio}_h^{\text{H1-13}}$ is the predicted approval ratio for the first half of 2013. Column 2 of Table 5 reports this zeroth stage. We find a strongly significant negative coefficient κ. Sectors with initially higher predicted

approval ratios experience larger increases in the penalty parameter when reserves fell. The sign of κ is consistent with the government having an ordering of which sectors to target with restrictions and being forced to go further down the list and restrict imports in initially-more-favored sectors when foreign reserves dwindled.

Table 5 contains the results of estimating our main specification using these predicted size penalties $\Delta\widehat{\phi}_{ht}$ as instruments (i.e. we regress price changes on quantity changes using the specification in (6) but at the half-year frequency and only within the DJAI period). We find coefficient estimates that are qualitatively consistent with those obtained with the between-policy regime variation. That is, positive shocks to the penalty parameter generate reductions in imported quantities and induce a negative relationship between quantities and prices.

Compared to the policy regime change analysis above, we are on firmer footing regarding endogeneity. Our identification arguments now rely on sectors with characteristics initially associated with high approval rates not subsequently being on differential trends that coincide with shocks to reserves (and recall reserves fell and then rose again rather than simply moving in one direction). It is also worth noting that since the sectors most affected by declines in reserves were the ones least effected in general by the DJAI policy, confounding trends that would bias the results in the policy regime analysis would bias results in the other direction in this analysis. Thus, it is reassuring that we find significant price increases with declines in quantity in both cases.

Table 5: The effects on changes in import prices and quantities during the policy

	(1)	(2)	(3)	(4)	(5)	(6)
			1^{st} stage	OLS	Reduced form	2^{nd} stage
	$\Delta\phi$	$\Delta\ln(q_I)$	$\Delta\ln(q_I)$	$\Delta\ln(p_I)$	$\Delta\ln(p_I)$	$\Delta\ln(p_I)$
$\Delta\ln(Reserves) \times \widehat{AR}^{\text{H1-13}}$	-0.015***		0.204***		-0.181***	
	(0.002)		(0.041)		(0.022)	
Predicted $\Delta\phi$		-9.336**				
		(3.525)				
$\Delta\ln(q_I)$				-0.236***		-0.883***
				(0.009)		(0.180)
Firm-product FE	Yes	Yes	Yes	Yes	Yes	Yes
Half-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	772,206	772,206	832,848	832,848	832,848	832,848
F-stat	41.4	7.0	24.6			

Notes: The sample is from the 1st half of 2014 to the 2nd half of 2015. The standard errors are one-way clustered by HS4-period and shown in parentheses. Asterisks indicate 10% (*), 5% (**), and 1% (***) significance. The predicted initial approval rate, $\widehat{AR}^{\text{H1-13}}$, is the fitted value from Table 2.

5.3 Bargaining Power

We now turn to the predictions concerning bargaining power. Recall that, according to the theory, the bargaining power weights should not affect the effect of the size penalty on import quantities. In contrast, the change in import price following a shock to the size penalty should be larger when the bargaining power of foreign suppliers is high.

To test these predictions, we proxy seller bargaining power with the number of importers which source from a particular HS8 product-country pair. We consider the following specification:

$$\Delta \ln y_{fit} = \alpha_0 \Delta\hat{\phi}_{ht} + \alpha_1 \left(\Delta\hat{\phi}_{ht} \times \ln F_b\right) + \mu_{fi} + \gamma_t + u_{fit}, \tag{10}$$

where $y \in \{q^I, p^I\}$, F_b is the number of firms that imported from a 8-digit product and origin pair b in $t = \text{Pre}$, f identifies firms, i identifies products (11-digit HS - origin - unit triplets), h identifies a 4-digit HS code, and

Table 6: Proxying seller power

	(1) $\Delta\ln(q_I)$	(2) $\Delta\ln(q_I)$	(3) $\Delta\ln(p_I)$	(4) $\Delta\ln(p_I)$
$\Delta\phi \times$ ln Buyers in HS8-Origin	0.033	0.048	0.156***	0.307***
	(0.09)	(0.16)	(0.05)	(0.09)
$\Delta\phi$	-2.212***	-2.705***	0.241	-0.322
	(0.32)	(0.51)	(0.17)	(0.33)
Firm-product FE	Yes	Yes	Yes	Yes
Period FE	Yes	Yes	Yes	Yes
N of Buyers	Any	>10	Any	>10
Observations	387,960	287,920	387,960	287,920
Clusters	10,995	9,756	10,995	9,756

Notes: The standard errors are clustered by firm and shown in parentheses. Asterisks indicate 10% (*), 5% (**), and 1% (***) significance. The variable "Buyers in HS8-Origin" is the number of firms that imported in a given HS8-Origin in the period before the policy (2011).

μ_{fi} and γ_t are firm-product and month fixed effects. The estimation results, in Table 6, lend support to the theory. The coefficient on the seller power proxy interaction is positive $\hat{\alpha}_1 > 0$ when we have prices on the left-hand side and it is zero $\hat{\alpha}_1 = 0$ when we have quantities on the left-hand side. In other words, our surprising finding of prices rises with quantity restrictions only holds for 8-digit product and origin pairs where there are many domestic importers and so the foreign supplier is likely to have substantial bargaining power.

6 Quantitative Evidence

TBA

7 Conclusion

TBA

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A Appendix

Contents

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A.4 Linking approvals to imports a.8

A.1 Additional Figures and Tables

Figure A.1: Our dataset compared to OECD

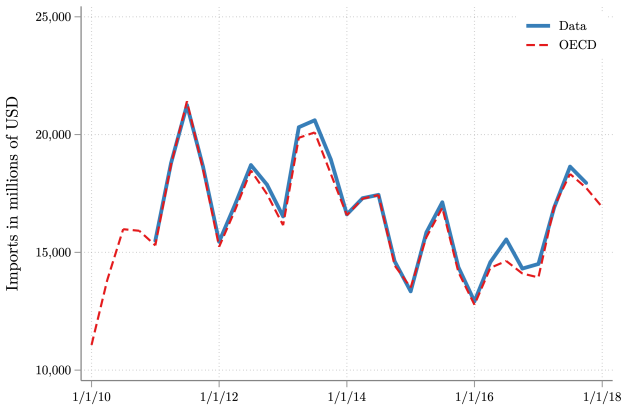


Table A.1: Distribution of estimates of ϕ and ϕ with different fixed effects.

(a) Firm and Month fixed effects

	Mean	1%	25%	Median	75%	99%	N
ϕ	0.043	-0.004	0.019	0.034	0.056	0.179	905
ϕ^{**}	0.047	0.007	0.022	0.037	0.059	0.193	774
ϕ_0	0.965	0.470	0.856	0.930	1.004	2.011	905
ϕ_0^{**}	0.967	0.480	0.857	0.931	1.004	2.011	903

(b) Saturated fixed effects
(Product, Firm, Month, Product-Firm, Product-Month, and Firm-Month)

	Mean	1%	25%	Median	75%	99%	N
ϕ	0.100	-0.146	0.036	0.073	0.133	0.673	731
ϕ^{**}	0.110	0.012	0.047	0.076	0.133	0.673	445
ϕ_0	1.337	-0.257	0.947	1.100	1.569	4.796	731
ϕ_0^{**}	1.374	0.575	0.969	1.118	1.541	4.796	593

Notes: Asterisks (**) indicate that we are keeping estimates significant at the 5% level. The standard errors are robust. When we saturate the model with fixed effects the number of estimates (N) falls because more categories (4-digit HS codes) do not have enough observations. The industry is the main 4-digit ISIC code of the firm.

Figure A.2: Size penalty (ϕ) estimates by year

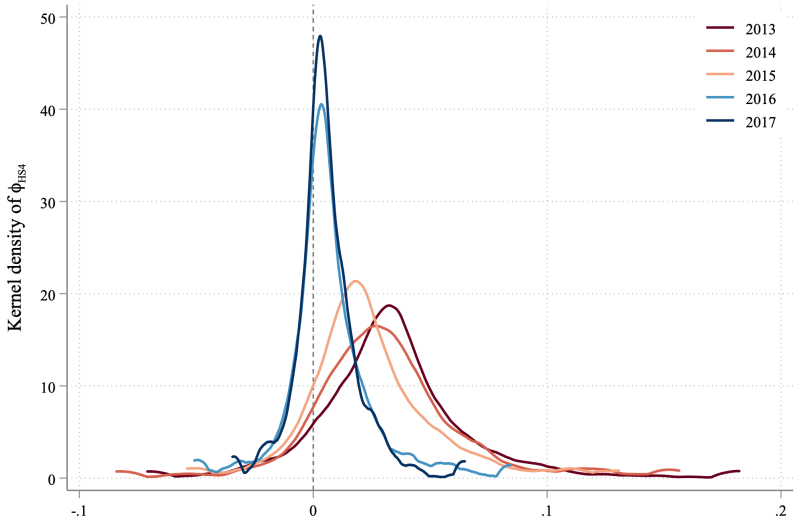


Figure A.3: Size penalty (ϕ) estimates with 95% confidence intervals

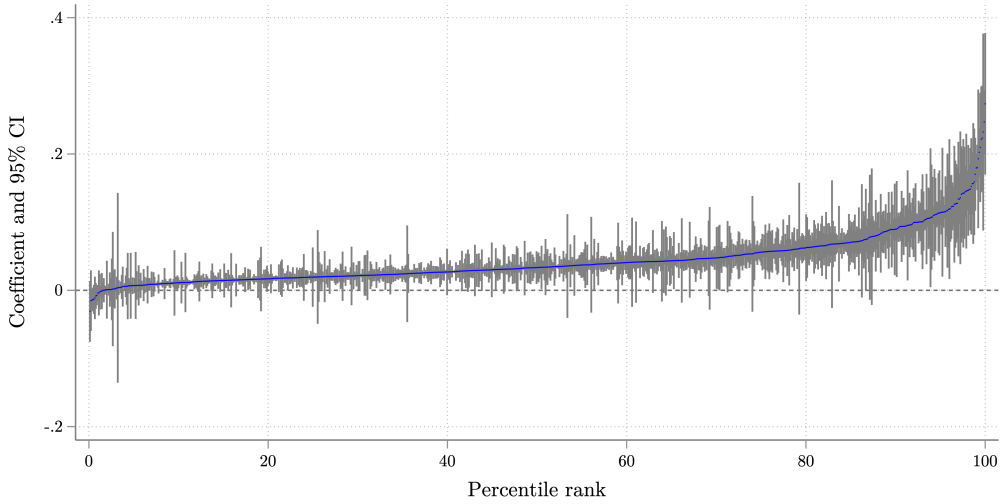
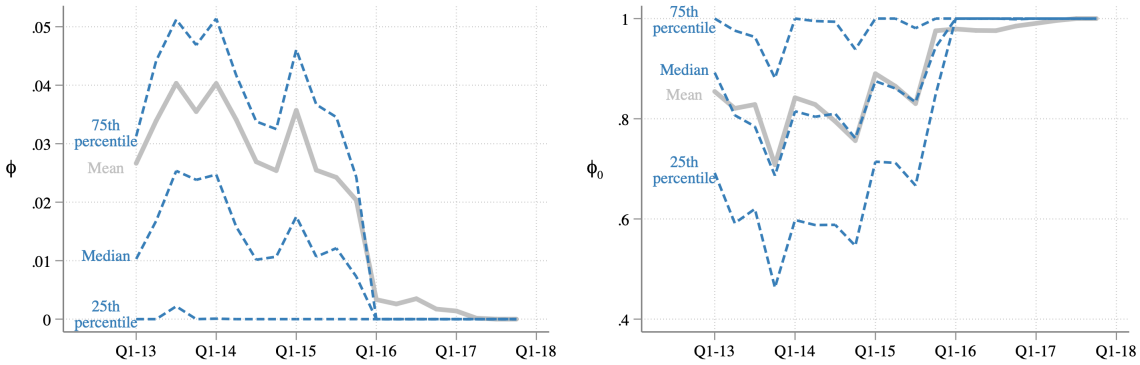
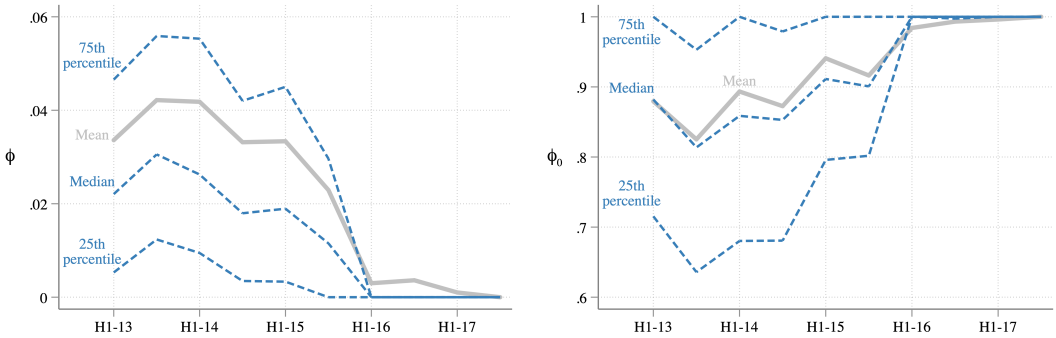


Figure A.4: Estimates of the size penalty (ϕ) and the approval level (ϕ_0) by quarter



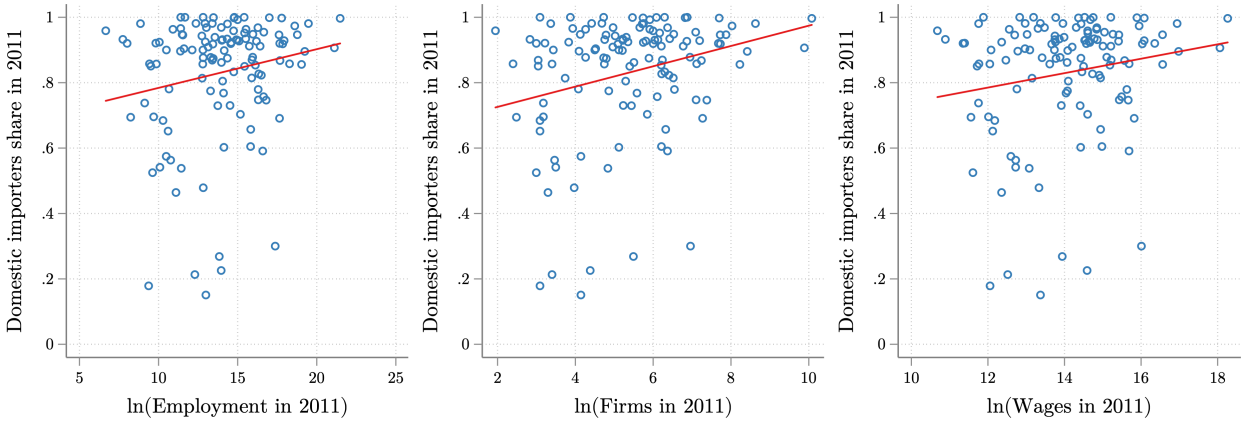
Notes: Moments of the estimates of ϕ (left panel) and ϕ_0 (right panel) at the level of the 4-digit HS by quarter-of-sample.

Figure A.5: Estimates of the size penalty (ϕ) and the approval level (ϕ_0) by half-year period



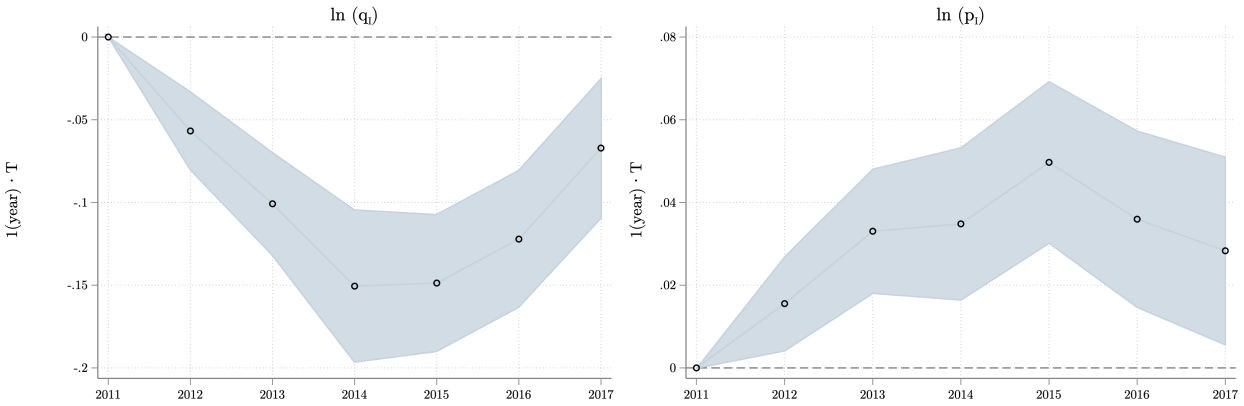
Notes: Moments of the estimates of ϕ (left panel) and ϕ_0 (right panel) at the level of the 4-digit HS by half-year period.

Figure A.6: Correlates of importer ownership by sector



Notes: Binned scatter plots and the line of best fit on the underlying data. Data at the level of a 4-digit ISIC (rev. 3) sector. Employment (number of jobs), number of firms, and total compensation come from the OEDE administrative data at the level of 4-digit ISIC sectors. To identify domestic importers we match the firm identifiers in our trade data to Orbis in order to get the global ultimate owner. Then we match 6-digit HS codes to 4-digit ISIC sectors and we compute the share of domestic over the total number of importing firms.

Figure A.7: Event-study with control group



Notes: *Notes:* Plot of the α_τ coefficients from estimating $y_{ift} = \alpha^\tau \mathbb{1}(t = \tau) \cdot T_h + \mu_{if} + \mu_{st} + \mu_t + error_{ift}$ with $t = 2011, \ldots, 2017$ and $\tau = 2012, \ldots, 2017$; i and f are product and firm identifiers, and h and s are the 4-digit HS code and the HS section of product i. The indicator T_h is 1 if ϕ_h is greater than the median, and 0 otherwise. The membership to a group is fixed in time. The LHS variable is imported quantities (q_I) in the left panel and import prices (p_I) in the right panel. The standard errors are clustered by firm.

Table A.2: Examples of product categories with high and low ϕ

ϕ quartile	4-digit HS	Description
1	2710	Petroleum oils, oils from bituminous minerals, not crude
	8708	Motor vehicle parts and accessories
	2711	Petroleum gases and other gaseous hydrocarbons
	2601	Iron ores and concentrates
	3004	Medicaments
2	8703	Motor cars and other motor vehicles
	3105	Fertilizers
	8408	Diesel or semi-diesel engines
	3808	Insecticides, rodenticides, fungicides, herbicides
	8407	Reciprocating or rotary internal combustion piston engines
3	8517	Line telephony or line telegraphy apparatus
	8704	Vehicles for the transport of goods
	8471	Automatic data processing machines
	8701	Tractors
	4011	New pneumatic tyres, of rubber
4	8802	Aircraft
	8502	Electric generating sets and rotary converters
	8429	Bulldozers, graders, levellers
	8428	Lifting, handling, loading or unloading machinery
	7207	Iron or non-alloy steel

Notes: Top sectors (4-digit HS codes) by value imported in 2011 for each quartile of the estimated size penalties (ϕ). The top quartile, number 4, has the highest ϕ's and thus is the most restricted.

Table A.3: Size penalty (ϕ) and importer ownership with controls

	ϕ in 2013-2015			
Domestic importer share in 2011	0.072***	0.067***	0.071***	0.069***
	(0.005)	(0.005)	(0.005)	(0.006)
ln(Imports in 2011)		-0.002***		0.000
		(0.001)		(0.001)
ln(Importers in 2011)			-0.005***	-0.004***
			(0.001)	(0.001)
ln(Exports in 2011)				-0.001***
				(0.000)
Constant	-0.009*	0.028*	0.014**	0.019
	(0.004)	(0.011)	(0.005)	(0.012)
Observations	886	886	886	868
F-stat	180.6	97.7	108.7	52.6

Notes: Robust standard errors shown in parentheses. Asterisks indicate 10% (*), 5% (**), and 1% (**) significance.

A.2 Description of the DJAI Procedure

The DJAI procedure was ordered by Argentina's Federal Public Revenue Administration (Administración Federal de Ingresos Públicos, AFIP) on January 5, 2012, and entered into force on February 1, 2012. Here we describe in more detail the steps involved in the procedure. The information in this section is adapted from the World Trade Organization Reports of the Panel.

First, importers were required to file an Advance Sworn Import Declaration (DJAI) through AFIP's electronic portal. The DJAI had to be submitted prior to the issuance of an order form, purchase order, or similar document issued to purchase items from abroad that were destined for local consumption.

The DJAI had to specify detailed information about the prospective import. That included value, quantity, origin, product code, measurement unit, and the name of the seller. We have all those fields in our data except for the name of the foreign seller. The product code is a 11-digit code that combines a 8-digit code from the NCM (Nomenclatura Común del Mercosur), and an extra 3-digit code specific to Argentina. The NCM is based on the Harmonized System (HS). For example, the 8-digit code 0901.11.10 contains coffee grains that are not roasted nor decaffeinated. The 11-digit codes split the beans further into arabica (0901.11.10.100) and robusta (0901.11.10.200).

A DJAI in "exit" (salida) status allowed the importation to proceed. Exit status on the DJAI was also necessary for obtaining authorization from the Central Bank of Argentina to make payments in foreign currency. If a DJAI in exit status was not used within 180 calendar days from initiation, the DJAI would be voided.

A DJAI would not enter "exit" status if a government agency entered "observations." Goods covered by a DJAI in "observed" (observada) status could not be imported into Argentina. A number of governmental agencies could review the information in the DJAI were entitled to enter "observations:"

- The Federal Public Revenue Administration (AFIP),
- The Secretariat of Trade (Secretaría de Comercio),
- The National Drugs, Food and Medical Technology Administration (Administración Nacional de Medicamentos, Alimentos y Tecnología Médica, ANMAT), and
- The Planning Secretariat for the Prevention of Drug Addiction and the Fight Against Drug Trafficking (Secretaría de Programación para la Prevención de la Drogadicción y la Lucha contra el Narcotráfico, SEDRONAR).

In general, an agency could enter an observation within 72 hours after a DJAI had been registered but the Secretariat of Trade had 15 working days to enter observations.

If a DJAI was in "observed" status, prospective importers could try to get the observation lifted. To do so, they were required to (a) identify the entity that entered the observation; (b) contact such agency in order to be informed of supplementary documents or information that must be provided; and, (c) provide the information required. If the agency lifted the observation within 180 calendar days from initiation, the DJAI would proceed to "exit" status. If the observation was not lifted within 180 calendar days from initiation, the DJAI would be voided.

A.3 Proofs of Theory Results

Proof of Claim 3.

Proof. Consider the first order condition associated to the expected surplus function in (2):

$$\left(\frac{1}{\sigma}aq^{-\frac{1}{\sigma}-1}+\frac{1}{\eta}zq^{\frac{1}{\eta}-1}\right)q^{1-\phi}=(1-\phi)\left(aq^{-\frac{1}{\sigma}}-zq^{\frac{1}{\eta}}\right)q^{-\phi}$$

Re-arranging immediately yields expression 3. Notice that $1-\frac{1}{\sigma}>\phi$ guarantees that $q_R>0$. Consider next the second order condition. The second derivative of the surplus function is given by:

$$q^{-\phi-1}\left(\left(1-\frac{1}{\sigma}-\phi\right)\left(-\frac{1}{\sigma}-\phi\right)aq^{-\frac{1}{\sigma}}-\left(\frac{1}{\eta}-\phi\right)\left(\frac{1}{\eta}+1-\phi\right)zq^{\frac{1}{\eta}}\right)$$

Evaluating this expression at the optimal quantity requested yields

$$q_R^{-\phi-1}a^{\frac{\frac{1}{\eta}}{\frac{1}{\eta}+\frac{1}{\sigma}}}z^{\frac{\frac{1}{\sigma}}{\frac{1}{\eta}+\frac{1}{\sigma}}}\left(-\left(1-\frac{1}{\sigma}-\phi\right)\left(\frac{1}{\sigma}+\phi\right)\left(\frac{1-\frac{1}{\sigma}-\phi}{1+\frac{1}{\eta}-\phi}\right)^{-\frac{\frac{1}{\sigma}}{\frac{1}{\eta}+\frac{1}{\sigma}}}-\left(\frac{1}{\eta}-\phi\right)\left(\frac{1}{\eta}-\phi+1\right)\left(\frac{1-\frac{1}{\sigma}-\phi}{1+\frac{1}{\eta}-\phi}\right)^{\frac{\frac{1}{\eta}}{\frac{1}{\eta}+\frac{1}{\sigma}}}\right)$$

This expression is negative if and only if

$$-\left(1-\frac{1}{\sigma}-\phi\right)\left(\frac{1}{\sigma}+\phi\right)<\left(\frac{1}{\eta}-\phi\right)\left(1+\frac{1}{\eta}-\phi\right)\left(\frac{1-\frac{1}{\sigma}-\phi}{1+\frac{1}{\eta}-\phi}\right) \tag{11}$$

If $1-\frac{1}{\sigma}-\phi>0$, then (11) boils down to $-\frac{1}{\sigma}<\frac{1}{\eta}$, which always holds. If $1-\frac{1}{\sigma}-\phi<0$, then (11) boils down to $-\frac{1}{\sigma}>\frac{1}{\eta}$, which never holds. □

Proof of Claim 3.

Proof. Note from 1 that ϕ does not enter into the $p^*(q)$ function. It follows that:

$$\frac{\partial}{\partial\phi}p^*=\frac{\partial p^*(q)}{\partial q}\frac{\partial q_R}{\partial\phi} \tag{12}$$

Claim 1 established that the latter term is negative. From 1, the former is:

$$\frac{\partial}{\partial q}p^*(q)=\frac{1}{\eta}\beta zq^{\frac{1}{\eta}-1}-\frac{1}{\sigma}(1-\beta)aq^{-\frac{1}{\sigma}-1}, \tag{13}$$

where q here denotes an arbitrary q. It follows that $\frac{\partial}{\partial\phi}p^* > 0$ when (13), evaluated at the optimal quantity requested q_R, is negative. That is, whenever:

$$\frac{\partial}{\partial q}p^*\left(q_R\right) < 0 \iff \frac{1}{\eta}\beta z q_R^{\frac{1}{\eta}-1} \leq \frac{1}{\sigma}\left(1-\beta\right)a q_R^{-\frac{1}{\sigma}-1}$$

Using (3), this condition can be written as

$$g\left(\phi\right) \equiv \frac{\sigma-1-\sigma\phi}{\eta+1-\eta\phi} \leq \frac{1-\beta}{\beta}$$

Note that $\phi < 1-1/\sigma$ for $q_R > 0$. In this range, $g\left(\phi\right)$ is decreasing and thus

$$\frac{\sigma-1-\sigma\phi}{\eta+1-\eta\phi} \leq \frac{\sigma-1}{\eta+1}$$

Thus:

$$\frac{\beta}{1-\beta} \leq \frac{\eta+1}{\sigma-1} \rightarrow p_I'\left(q_I\left(\phi\right)\right) < 0$$

Alternatively:

$$\frac{\beta}{1-\beta} \leq \frac{\eta+1-\eta\phi}{\sigma-1-\sigma\phi}$$

The RHS is an increasing function of ϕ, so that

$$\frac{\eta}{\sigma-1} \leq \frac{\eta+1}{\sigma-1} \leq \frac{\eta+1-\eta\phi}{\sigma-1-\sigma\phi}$$

$$\frac{\beta}{1-\beta} < \frac{\eta+1-\eta\phi}{\sigma-1-\sigma\phi} = \frac{\frac{\eta+1}{\phi}-\eta}{\frac{\sigma-1}{\phi}-\sigma} \leq \frac{\eta}{\sigma}$$

Thus

$$\frac{\beta}{1-\beta} \leq \frac{\eta}{\sigma}$$

is a sufficient condition for $p_I\left(\phi\right)$ to increase with ϕ. □

Proof of Claim (4).

Proof. Take a derivative of (12) with respect to β :

$$\frac{\partial^2}{\partial\beta\partial\phi}p^* = \frac{\partial^2 p^*\left(q\right)}{\partial\beta\partial q}\frac{\partial q_R}{\partial\phi} + \frac{\partial p^*\left(q\right)}{\partial q}\frac{\partial^2 q_R}{\partial\beta\partial\phi}$$

Note from (3) that q_R doest no depend on β and hence

$$\frac{\partial^2 q_R}{\partial\beta\partial\phi} = 0$$

If follows that:

$$\frac{\partial^2}{\partial\beta\partial\phi}p^* = \frac{\partial^2 p^*\left(q\right)}{\partial\beta\partial q}\frac{\partial q_R}{\partial\phi}$$

Recall that the second term in the right hand size of the above expression is negative (Claim 1). Relying on

(13), the first term is given by:

$$\frac{\partial^2 p^*}{\partial\beta\partial q} = z\frac{1}{\eta}q^{\frac{1}{\eta}-1} + \frac{1}{\sigma}aq^{-\frac{1}{\sigma}-1} > 0$$

□

A.4 Linking approvals to imports

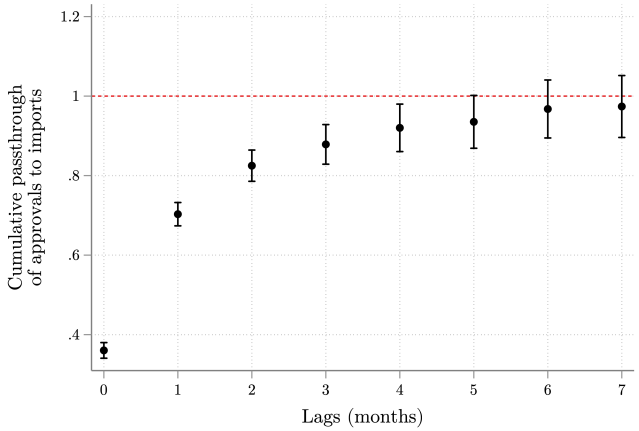
In this section we argue, in two different ways, that approvals were converted to imports. First, we run a regression of changes in imported quantities on lags of changes in approved quantities. Second, we estimate an equation of log imported quantities on log approved quantities but instrumenting for log approved quantities.

The first results (Figure A.8) show a passthrough regression of the form

$$\Delta\ln(q^I_{fit}) = \alpha + \sum_{k=0}^{7}\beta^k\Delta\ln(q^A_{fit-k}) + e_{fit}, \tag{14}$$

where t is a month-of-sample, i is a product, and f is a firm. The figure plots the cumulative partial sum of the β^k coefficients and 95% confidence intervals. An increase in approvals of 10% is associated with an increase in imports slightly under 4% on that same month, but also to subsequent increases in imports so that after five months the accumulated increase is 10%. That is consistent with the institutional characteristics that we leared about the DJAI, since a requested import could only be converted to actual imports within 180 days of the initiation of the request.

Figure A.8: Passthrough from quantity approved to imported



Notes: Cumulative partial sums of the β^k coefficients from the estimation of equation 14 and their 95% confidence intervals. The value for zero months on the horizontal axis is β^0, the one for one month is $\beta^0+\beta^1$, and so on. The standard errors are clustered by firm-product. The data is monthly aggregates in 2014.

Second, we show that plausibly exogenous increases in the approved quantity translated almost one-to-one to increases in imports. We construct an instrument that interacts changes in the exchange rate gap during 2014-2015 with mean quantities approved during 2013:

$$\frac{G_t}{\overline{G}} \times \ln(\overline{Q}^A_{fi}) \tag{15}$$

where G_t is the ratio of monthly formal to informal exchange rate in ARS per USD, t is a month-of-sample, $\overline{G} \equiv \frac{1}{12}\sum_{k=\text{Jan-13}}^{\text{Dec-13}} G_k$ is the mean gap in 2013, and $\overline{Q}^A_{fi} \equiv \frac{1}{12}\sum_{k=\text{Jan-13}}^{\text{Dec-13}} q^A_{fik}$ is the average approved quantity

during 2013 for product-firm fi. The economic rationale is that when the exchange rate gap is larger the government will have more pressure to restrict imports, and would do so by restricting more those firms and products that it has revealed it prefers to restrict in the early stages of the policy. We use this instrument to obtain a two-stage least squares estimate of the equation in levels of approvals on imports:

$$\ln(q^I_{fit}) = \alpha + \beta \ln(q^A_{fit}) + e_{fit}. \tag{16}$$

Table A.4 shows the results. The naive OLS of log quantites approved on imported is positive and significant at 0.376, meaning that a 10% increase in approvals is associated with a 3.76% increase in imports. The IV estimate is higher and closer to one at 0.854. In fact, in the last row of the last column we see that the 95% confidence interval associated with the IV estimate includes 1. That is, we cannot reject the null hypothesis that a 10% increase in approvals leads to a 10% increase in imports.

Table A.4: Imported quantities on instrumented approved quantities

	OLS	1^{st} stage	Reduced form	2^{nd} stage
	$\ln(q_I)$	$\ln(q_A)$	$\ln(q_I)$	$\ln(q_A)$
$\ln(q_A)$	0.376^{***}			0.854^{***}
	(0.03)			(0.23)
$\frac{G}{\bar{G}}\ln(\bar{Q}^A)$		0.050^{***}	0.042^{**}	
		(0.01)	(0.01)	
Firm-product FE	Yes	Yes	Yes	Yes
Month-of-sample FE	Yes	Yes	Yes	Yes
Observations	994,998	994,998	994,998	994,998
Importers	7,558	7,558	7,558	7,558
F-stat		13.5		
95% CI				[0.40,1.30]

Notes: Two-stage least squares estimates of Equation 16 using the instrument defined in 15. The standard errors are clustered by firm and shown in parentheses. Asterisks indicate 10% (*), 5% (**), and 1% (**) significance. The last row of the last column shows the 95% confidence interval associated with the IV estimate, which includes 1.