Tariff Rate Uncertainty and the Structure of Supply Chains*

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Abstract

We develop a model of international procurement and show that increases in the probability of trade war inhibit domestic buyers' use of long-term relationships to ensure the provision of high-quality inputs. We introduce a method for identifying such "Japanese" importing in transaction-level trade data, and find that a policy change that lowered the likelihood of trade war fostered a relative shift towards "Japanese" procurement among U.S. importers and Chinese exporters. Embedding our model in a standard multi-country general equilibrium model of Ricardian comparative advantage, we show via quantitative simulations that such shifts can raise welfare appreciably by allowing buyers to avoid costly inspection. (JEL Codes: F13, F14, F15, F23) (Keywords: Supply Chain, Uncertainty, Trade War, Procurement)

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1 Introduction

Between 1990 and 2008, the rapid expansion of global value chains promoted a substantial increase in world trade and unprecedented convergence in rich and poor country incomes (WorldBank, 2019). Since the global financial crisis, however, increasing restrictions have impeded trade and dampened firms' enthusiasm for expanding supply chains internationally, threatening developing countries' export-led growth and reducing welfare in a variety of ways, as documented in Bown, Conconi, Erbahar, and Trimarchi (2020), Amiti, Redding, and Weinstein (2019), Fajgelbaum, Goldberg, Kennedy, and Khandelwal (2019), Flaaen and Pierce (2019), and Flaaen, Hortaçsu, and Tintelnot (2020). In this paper, we focus on another implication of these shifts in policy, demonstrating that increases in the probability that trade might be restricted can cause supply chains to be re-shuffled and welfare to decrease by raising firms' procurement costs.

In the first part of the paper, we develop a model of global sourcing that builds upon the partial-equilibrium framework for domestic supply chains introduced by Taylor and Wiggins (1997), where buyers choose between two simple procurement systems to minimize the cost of guaranteeing high-quality inputs from sellers.[1] Under the "Japanese" system, buyers motivate a single seller to maintain high input quality by committing to smaller, more frequent purchases at a price above cost over a long-term relationship. In the opposing "American" system, buyers choose larger, less frequent purchases at lower prices from a parade of different sellers in the spot market. In this system, costly inspections and enforceable contracts deter sellers from cheating on product quality. Lower inspection costs favor the "American" system, while policy environments that promote the formation of long-term relationships favor the "Japanese" system.[2]

We extend Taylor and Wiggins (1997) to *international* procurement by linking domestic importers' ability to maintain long-term relationships with foreign sellers to changes in the arrival rate of a trade war. In equilibrium, each buyer procures and distributes its product using the procurement system that minimizes costs. We show that increases in the probability of a trade war reduce the likelihood that buyers choose "Japanese" procurement because it shortens the expected length of buyer-seller relationships, thereby raising the premia buyers must pay sellers to incentivize high quality.

In the second part of the paper, we employ confidential transaction-level U.S. import data from the U.S. Census Bureau to investigate the prevalence of "Japanese" sourcing among U.S. importers and examine the implications of the model. These data are well-suited for our inquiry as they record both the number of foreign exporters with which U.S. importers trade, as well as quantities and therefore unit values (prices) associated with each shipment. Guided by our model, we classify U.S. importing firms as either "Japanese" or "American" based on the number of foreign suppliers from which they purchase a particular product from each country over the sample period, 1992 to 2016.

[1] More broadly our paper is related to the literature on repeated games (e.g., Green and Porter (1984), Abreu, Pearce, and Stacchetti (1990)), and in particular to the literature on repeated games with incomplete information, see Kandori (2002). We focus on empirically testing one specific framework.

[2] The global movement towards "Japanese" procurement is documented in a series of studies in the management and economic literatures, e.g., O'Neal (1989), Heide and John (1990), Lyons, Krachenberg, and Henke Jr. (1990), Dyer and Ouchi (1993), Han, Wilson, and Dant (1993), Helper and Sako (1995), and Liker and Choi (2004).

We find that "Japanese" importing during the late 1990s is most prevalent from Japan and Mexico, and that it increases most substantially in the early 2000s for Brazil, China and Mexico. We then show, consistent with the model, that buyers with a lower ratio of sellers to shipments – i.e., those that are more "Japanese" – do indeed receive smaller, more frequent shipments at a higher price than "American" buyers of the same product. To our knowledge, these results provide the first systematic empirical evidence identifying and rationalizing "Japanese" and "American" procurement patterns among buyers and sellers in a large-scale dataset.

In the third part of the paper, we compare U.S. importing relationships before and after the 2001 US extension of Permanent Normal Trade Relations (PNTR) towards China, which dramatically reduced the possibility of a trade war between the two countries. Exploiting the fact that the impact of this change in policy varied substantially across products, we employ a triple difference-in-differences specification that asks whether U.S. importers' procurement patterns change after the policy is implemented (first difference), for imports from China relative to other countries (second difference), in products with greater relative exposure to PNTR (third difference).[3] Results are consistent with the model: we find that shipments of more-exposed products from China do indeed become relatively smaller, more frequent, and more expensive after 2001, indicating a switch towards "Japanese" procurement. Coefficient estimates suggest that a one standard deviation increase in exposure to PNTR is associated with a relative decline in average shipment quantity and increase in frequency of 4.5 percent, and a relative increase in average shipment price of 2.1 percent.

In the final section of the paper, we embed our model of procurement in an extension of Eaton and Kortum (2002)'s multi-country, general equilibrium model of Ricardian comparative advantage and analyze the change in trade flows and the welfare associated with an increase in the probability of trade conflict. In the resulting framework, trading patterns are governed by bilateral probabilities of trade war, in addition to standard cross-country differences in productivity, as they alter the relative attractiveness of the "Japanese" system across trading partners. Quantitative simulations of the model reveal that an increase in the probability of trade war that is sufficiently large to eliminate "Japanese"-style procurement increases domestic sourcing at the expense of imports. The associated rise in final goods prices due to higher procurement costs lowers U.S. welfare by more than 3 percentage points. The relevance of this scenario is underscored by recent, sharp increases in policy uncertainty in important bilateral trading relationships including US-China and UK-EU trade.

Our analysis makes contributions to several literatures. First, we add to the growing body of research on trade wars and trade policy uncertainty (Ossa, 2014; Handley, 2014; Handley and Limão, 2017; Alessandria, Khan, and Khederlarian, 2019) by identifying a new channel – procurement systems – through which changes in the probability of trade war can influence trade patterns and welfare. Our results with respect to changes in unit values as firms switch between procurement systems also highlight a novel source of price variation in response to changes in trade policy that goes beyond the quality premiums and markups currently studied (Goldberg and Pavcnik, 2016; De Loecker and Goldberg, 2014; Antoniades, 2015; De Loecker, Goldberg, Khandelwal, and Pavcnik, 2016; Manova

[3] In our model, buyer and seller trade a single product, so the probability of a trade war and the probability the buyer-seller relationship ends are the same. Our empirical analysis, on the other hand, examines firms trading a wide range of products subject to varying potential increases in tariffs prior to the granting of PNTR to China.

and Yu, 2017; Kugler and Verhoogen, 2012; Verhoogen, 2008; Khandelwal, 2010).

Second, we contribute to greater understanding of the organization and structure of global value chains (Antras and Chor, 2021; Antràs and Chor, 2018; Antràs, Fort, and Tintelnot, 2017), as well as a larger literature on incomplete contracts, imperfect contract enforcement and information asymmetries that primarily focuses on buyer-seller relationships within countries (Antràs, 2003, 2005; Antràs and Helpman, 2008; Feenstra and Hanson, 2005; Fisman and Wang, 2010; Grossman and Helpman, 2004; Spencer, 2005).[4] In contrast to much of the research in this area, we consider "Japanese sourcing" rather than firm integration as a solution to firms' quality-control problem. This path is particularly relevant for understanding sourcing in settings where integration is not possible, for example, in China, where foreign firms face numerous formal and informal restrictions regarding ownership of domestic assets. Another alternative to integration pursued in the literature is relational contracting, i.e., repeated transactions between buyers and sellers under an informal agreement (Defever, Fischer, and Suedekum, 2016; Kukharskyy, 2016).[5] In that literature, the pattern of trade between buyers and sellers is governed by idiosyncratic time preferences. Here, by contrast, discount rates are common and firms choose between procurement systems based on inspection costs and the probability of a trade war. As a result, our model links shipment patterns to trade policies in a manner amenable to empirical inquiry using transaction-level trade data.[6]

Third, our analysis complements recent empirical studies using shipment-level trade data to study trade frictions and inventory costs. Existing research shows that high fixed per-shipment trade costs reduce shipping frequency, thereby raising inventories in a manner that can have substantial impact on how firms adjust to trade shocks (Kropf and Sauré, 2014; Hornok and Koren, 2015b,a; Békés, Fontagné, Muraközy, and Vicard, 2017; Alessandria, Kaboski, and Midrigan, 2010, 2011). Here, we show that trade policy uncertainty can be an important barrier to firms' efforts to reduce inventory costs, as it promotes use of the "American" system. An interesting question for future research is whether minimizing such costs provides an additional motivation to lobby for (more dependable) free trade.

Finally, we contribute to the burgeoning literature using general equilibrium models to evaluate changes in trade policy by extending Eaton and Kortum (2002) in two directions. First, as mentioned above, in our model, product prices depend on the probability of a trade war as well as supplier productivity. Second, we allow for increasing returns to scale among sellers as, in our setting, sellers incur fixed logistics costs in addition to variable costs in producing each shipment. We account for the resulting declining average cost curves by assuming sellers compete in a contestable market (Baumol, Panzar, and Willig, 1982), a natural extension to price competition in the presence of

[4] Recent research on domestic procurement includes Tadelis and Zettelmeyer (2015), Cicala (2015), Bajari, Houghton, and Tadelis (2014), Macchiavello and Morjaria (2020), and Kukharskyy and Pflüger (2010).

[5] Macchiavello and Morjaria (2015), for example, examine the value of reputation for reliability among Kenyan rose exporters.

[6] In contrast to heterogeneous-firms models of trade and investment (Melitz, 2003; Bustos, 2011), the fixed and variable costs of trade in our setting are endogenous to firms' choice of procurement system. Our results also relate to a broader set of empirical papers analyzing trade flows among importer-exporter relationships (Monarch and Schmidt-Eisenlohr, 2017; Huneeus, 2018; Kikkawa, Magerman, and Dhyne, 2019; Heise, 2019; Alviarez, Fioretti, Kikkawa, and Morlacco, 2021; Flaaen, Haberkorn, Lewis, Monken, Pierce, Rhodes, and Yi, 2019), though not in the context of procurement system choice.

economies of scale (Tirole, 1988). We use the model to quantify how a reduction in buyers' ability to form "Japanese" relationships reduces welfare by raising prices, similar to a negative productivity shock. While we develop the model in the context of changes in the probability of trade peace, the underlying manner by which trade regimes affect productivity applies more broadly to any factor that might undermine sellers' beliefs about the viability of establishing long-term relationships with buyers, e.g., uncertainty over the arrival of shipments due to corruption, pandemics, bad weather, or disruptions at ports.

The remainder of this paper proceeds as follows. Section 2 outlines our theoretical model. Section 3 describes the data, and Section 4 presents empirical evidence for Japanese style procurement. Section 5 analyzes the effect of a change in trade policy uncertainty. Sections 6 and 7 embed our partial equilibrium model into Eaton and Kortum (2002) and perform quantitative simulations. Section 8 concludes. An online appendix contains additional explanatory detail and results.

2 Theoretical Model

Incomplete contracts, information asymmetries and contract enforcement are common problems associated with firms' purchases of intermediate inputs. While firm integration is one means of addressing these issues (Antràs, 2003, 2005; Antràs and Helpman, 2008), it may not be a viable solution if cross-border integration involves substantial costs.[7] As a result, we focus on an alternate, arm's-length solution to the quality control problem pioneered by Japanese firms such as Toyota.

Our starting point is the setup introduced by Taylor and Wiggins (1997), in which a buyer seeks to obtain inputs from a supplier whose effort is unobservable. This type of problem falls into the class of repeated games with incomplete information (see, e.g., Kandori (2002), Mailath and Samuelson (2006)). The framework by Taylor and Wiggins (1997) is particularly suitable to our context because it delivers as optimal contract solving the repeated game one of two simple procurement systems, which have been described anecdotally in the literature as capturing the main dimensions by which firms' procurement strategies differ, e.g., Helper and Sako (1995). Under the "American" system, buyers use competitive bidding to select the lowest-cost supplier for each shipment of inputs, and use the threat of inspection to deter provision of low-quality goods. Under the "Japanese" system, by contrast, buyers offer sellers incentive premia over long-term relationships to ensure inputs of sufficient quality. The two procurement systems map into observable order patterns which we can examine in transaction-level trade data to classify firms' procurement strategies. We extend Taylor and Wiggins (1997) to an international trade context by linking incentive premia to the arrival rate of a trade war (this section), and by endogenizing final demand (Section 6). These extensions allow us analyze the effect of a change in the probability of trade peace on a host of outcomes, including trade flows and welfare.

[7] China, for example, requires foreign ventures to include a domestic partner, while the United States (and other developed countries) mandate national security reviews of foreign acquisitions in certain sectors.

2.1 The Procurement Problem

2.1.1 The Seller's Problem

There is a single country populated by a continuum of homogeneous sellers able to produce the same good.[8] To complete a production run (i.e., produce one shipment) a seller hires labor l at wage $w = 1$ to produce and delivers output $x = \frac{\Upsilon}{\theta} l$, where Υ is a seller's productivity and θ represents the seller's product's level of quality. The unit input requirement, $\frac{\theta}{\Upsilon}$, allows for variation in quality, giving rise to a "quality control" problem.[9] Sellers choose between discrete quality levels, $\theta \in \{\underline{\theta}, \overline{\theta}\}$, where lower quality is less costly to produce. To complete the shipment, the seller absorbs f units of labor for per-shipment specific logistics services, including transport costs.[10] The seller's total costs for each production and delivery cycle are therefore $x\frac{\theta}{\Upsilon} + f$.

2.1.2 The Buyer's Procurement Choices

There exist multiple homogeneous buyers that are willing to procure a seller's output and distribute it downstream in the consumer market. These buyers compete in a contestable market, described in more detail below. Conditional on desired quality, $\overline{\theta}$, consumer demand arrives continuously. Let t denote continuous time and consider time periods $\Delta t = \int_0^1 1dt = 1$, e.g., 1 year. To supply the consumer market over one time period, a buyer procures total quantity, q, in a series of discrete, equally sized, symmetric shipments of size x. We take q as fixed in this section, but solve for it in equilibrium in Section 6. Consequently, there are q/x shipments during each period. Figure 1 summarizes the shipment and consumption pattern visually. If quality is less than desirable, no downstream consumer demand arrives for the product and buyers must dispose of the obsolete shipment without recompense. To avoid these losses, the buyer seeks to ensure the provision of high-quality inputs using either an "American" (A) or a "Japanese" (J) procurement system.

In the "American" (A) system, buyers pay fixed cost m_A to inspect each shipment's quality before delivery. We assume inspections reveal quality with certainty.[11] Therefore, if buyers inspect, sellers cannot gain by cheating on product quality. Consequently, buyers know that inspections guarantee product quality. We assume that buyers have all the bargaining power. As a result, given an order of size x_A placed with a seller, the buyer sets the per shipment price $v_A(x_A, \overline{\theta})/x_A$ to allow the seller to break even and participate, where

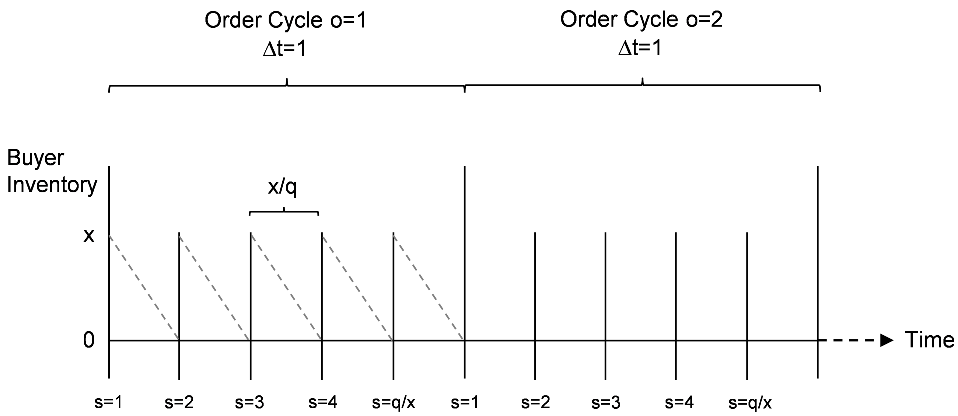
[8] We extend the model to multiple products and sellers in multiple countries in Section 6 below.

[9] See, for example, "Poorly Made," The Economist, May 14th, 2009.

[10] While a large international trade literature models the cost of shipping between origin and destination as "iceberg", i.e., a fractional loss of the shipped good which rises with distance, recent evidence supports per-unit and per-shipment specific delivery costs (Hummels and Skiba, 2004; Martin, 2012; Kropf and Sauré, 2014; Hornok and Koren, 2015b,a). We note that our theory could incorporate an iceberg cost in addition to fixed costs without changing the conclusions; however, it is important that at least some component of costs is fixed to generate discrete shipments.

[11] Taylor and Wiggins (1997) allow for a more general inspection pattern and show that optimal inspection frequency under the American system is a function of shipment size and quality. This generalization does not affect conclusions below related to per-shipment costs. Our simplifying assumption allows us to derive the properties of the model via implicit function techniques, as opposed to relying on the near-linearity of the problem when r/q is small as in Taylor and Wiggins (1997). We cannot rely on a small discount factor since in our model discounting is additionally affected by the arrival rate of trade wars.

Figure 1: Timing



Notes: The total quantity shipped over an order cycle is q. Order cycles repeat indefinitely and are indexed by $o = \{1, 2, ...\}$. There are $s = \{1, 2, ..., q/x\}$ shipments during an order cycle, each arriving every x/q units of time apart.

$$v_A(x_A, \overline{\theta}) = f + \frac{\overline{\theta}}{\Upsilon} x_A. \tag{1}$$

Due to the fixed cost, the buyers' average procurement costs are decreasing in order size, and therefore each buyer optimally places each order with a single seller. Since the sellers are homogeneous and all willing to supply at the same price, we assume that for a given buyer the winning seller is chosen randomly for each order. Inclusive of inspection costs, the buyer's total procurement expense equals $v_A(x_A, \theta) + m_A$. The form of this procurement cost is similar to those appearing elsewhere in the literature (Kropf and Sauré, 2014; Hornok and Koren, 2015b,a), where exogenous per-shipment fees such as f and m_A capture administrative barriers.

"Japanese" (J) procurement motivates the production of high quality via payments of a premium over a seller's cost over the course of a long-term relationship. A seller chooses to ship high quality if a long-run relationship with the buyer is of sufficient value, and a contribution of this paper is to show how this value depends upon the stability of trade policy between seller and buyer countries. We assume trade policy shocks that break buyer-seller relationships, e.g., an escalation of the tariff on the product to a prohibitive level, arrive at a constant rate, ρ.[12] In that case, relationships break before time t with probability $F(t) = 1 - e^{-\rho t}$, implying survival over a shipment cycle with probability $e^{-\frac{\rho x}{q}}$.[13] We note that while our focus is on trade policy, there are many other shocks which might have similar effects. For example, natural disasters may affect suppliers' ability to reliably fulfill their implicit agreements with buyers in foreign countries (Boehm, Flaaen, and Pandalai-Nayar, 2019).

[12] Ossa (2014), for example, estimates that the optimal tariffs countries might set in the event of a trade war are substantial, averaging 63 percent worldwide.

[13] For proof, see Wooldridge (2002), page 688. Note that the model considers trade in a single product. An alternate interpretation of ρ related to our data analysis below is that it reflects the overall probability of a buyer-seller relationship becoming unprofitable due to a change in trade policy. That is, ρ reflects both the probability of a trade war (which is the same for all products) and the magnitude of the subsequent rise in tariffs (which might vary across products). The probability of relationship breakup is rising in both of these factors.

For our purposes, $e^{-\frac{\rho x}{q}} < 1$ implies that firms are uncertain about whether future trade policy facilitates relationships: for a given shipping cycle $\frac{x_J}{q}$, a greater arrival rate of trade wars, ρ, increases the separation probability. Let r be the per-period interest rate and $v_J(x_J, \theta)$ be the payment the buyer sets under the "Japanese" system for each shipment. With continuous compounding, the expected discounted value of the relationship over all future shipment cycles is then $\frac{v_J(x_J,\bar{\theta})}{1-e^{-(r+\rho)x_s/q}}$.[14] Note that here, in contrast to the "American" system, a buyer procures each order from the same seller.

To guarantee desired quality, the buyer must set a per-shipment payment such that the seller's net present value of the continued relationship exceeds the one-time profit from cheating by supplying inferior quality,

$$\frac{v_J(x_J, \bar{\theta}) - f - \frac{\bar{\theta}}{\Upsilon} x_J}{1 - e^{-(r+\rho)x_J/q}} \geq v_J(x_J, \bar{\theta}) - f - \frac{\underline{\theta}}{\Upsilon} x_J. \tag{2}$$

In this expression, we assume that if the seller provides low quality, the buyer does not find out about it until after the shipment is received and the payment is made, and that the seller delivering low quality is excluded from the market forever. Here, too, we assume buyers have all the bargaining power. Solving (2), buyers under the "Japanese" system set the per-shipment payment to be

$$v_J(x_J, \bar{\theta}) = f + \bar{\theta} \frac{1}{\Upsilon} x_J + \left[e^{(r+\rho)x_J/q} - 1 \right] (\bar{\theta} - \underline{\theta}) \frac{1}{\Upsilon} x_J. \tag{3}$$

In the "Japanese" system buyers pay the per-unit premium $\left[e^{(r+\rho)x_J/q} - 1 \right] (\bar{\theta} - \underline{\theta}) \frac{1}{\Upsilon}$ to induce the seller to provide high quality.

A key feature of the "Japanese" system, therefore, is that more stable trade relationships (i.e., lower separation rates ρ) and smaller shipments, x_J, sent more frequently (which increases the present discounted value of payments) reduce the premium necessary to guarantee desired quality.[15] As a result, compared to a setting without incentive problems (e.g., an integrated firm), "Japanese" procurement has higher variable costs while American procurement system has higher fixed costs.

Buyers choose between the "American" and "Japanese" system by comparing long-term expected revenues and costs. We assume that a trade war causes buyers to exit irrespective of the system they choose, as they lose access to the suppliers of their good.[16] At a given market price p, long-term expected profits in the two procurement systems are given by

$$\pi_s^b = \left[\int_0^{x_s/q} e^{-rt} pq \, dt - v_s(x_s, \bar{\theta}) - m_s \right] / \left[1 - e^{-(r+\rho)x_s/q} \right] \quad \text{s} \in \{\text{J,A}\} \tag{4}$$

[14]The discount rate over a shipping cycle with associated continuous discount factor is $\lim_{N \to \infty} \left(\frac{1}{1+\frac{rx}{q}/N} \right)^N = e^{-\frac{rx}{q}}$.

[15]An alternative approach to incorporating trade policy uncertainty would be to multiply the discount factor by an exogenous probability of trade peace $(1-\rho)$. However, a drawback of that approach is that the probability of relationship separation over a given time period is dependent of the number of shipments made. In our formulation, the likelihood of separation is independent of shipment frequency.

[16]In the model with multiple sellers discussed below, an alternate assumption is that buyers switch to a seller from another country in the event of a trade war. Given that buyer profits are zero in equilibrium, however, this assumption is equivalent to assuming that the buyer exits.

where discounted revenues at the beginning of each shipment cycle are $\int_0^{x_s/q} e^{-rt} pq\, dt$ and $m_J = 0$.

2.2 Market Equilibrium and Optimal Procurement Choice

We now study the optimal procurement pattern and system for a given fixed order quantity q. We assume in this section that buyers' profits are zero, and show in Section 6 that our contestable markets setup delivers zero profits as an equilibrium outcome with endogenous q. Setting profits equal to zero implies that the market price must equal average procurement and distribution costs, $AC_s\,(x_s, q)$, and hence we obtain the following zero-profit conditions conditional on the procurement system,

$$p_s = AC_s\,(x_s, q) = \left(\frac{r}{q}\right) \frac{v_s(x_s, \bar{\theta}) + m_s}{\left[1 - e^{-rx_s/q}\right]} \quad \text{s}\in\{\text{J,A}\}, \tag{5}$$

where $m_J = 0$. Buyers choose a shipment size to minimize average procurement costs within each procurement system. Taking first order conditions (FOC_s) for each system and setting them to zero we obtain,

$$\frac{v_s'(x_s, \bar{\theta})}{1 - e^{-rx_s/q}} = \frac{\left[v_s(x_s, \bar{\theta}) + m_s\right] \frac{r}{q} e^{-rx_s/q}}{\left(1 - e^{-rx_s/q}\right)^2} \quad \text{s}\in\{\text{J,A}\}. \tag{6}$$

This expression implicitly determines the optimal shipment size, x_s^*. The left hand side represents the discounted value of higher costs associated with a small increase in order size. The right hand side measures the savings from an increased discount factor due to spacing these larger orders further apart in time. Trading off these costs and benefits, the firm optimally procures x_s^* to minimize average expected purchasing costs. We present all proofs in Appendix A, and show in Appendix A.1.1 that an interior solution to the first order condition is a unique cost minimizer for $0 < rx/q < 1$ under both procurement systems.

Conditional on procurement quantity q and parameter values, the buyer compares average procurement costs evaluated at the optimum, $AC_s\,(x_s^*, q)$, across systems to determine the cost-minimizing procurement system. Implicit function techniques provide intuition for this comparison in the presence of non-linearity. With no variation in quality, i.e., for $\bar{\theta} - \underline{\theta} = 0$, and with $m_A = 0$, there is no incentive problem and costs in both systems are identical. Compared to this benchmark case, differentiating equation (5) under the "Japanese" system with respect to $\underline{\theta}$ and ρ, respectively, and noting that, by the envelope theorem, the indirect effect coming from the resulting change in x_J is zero, we find that average procurement costs in the "Japanese" system increase with the arrival rate of trade wars, ρ, and with the range of potential qualities, $\bar{\theta} - \underline{\theta}$, due to the greater incentive premia they necessitate, $\frac{\partial AC_J\left(x_J^*, q\right)}{\partial \underline{\theta}} \leq 0$ and $\frac{\partial AC_J\left(x_J^*, q\right)}{\partial \rho} \geq 0$.[17] In the "American" system, differentiating (5) with respect to m shows that average costs increase with inspection costs m. Importantly, as $m \to \infty$, we have $AC_A\,(x_A^*, q) \to \infty$ because average costs grow without bound, $\frac{\partial AC_A\left(x_A^*, q\right)}{\partial m} = \frac{1}{1 - e^{-\frac{rx_A^*}{q}}} > 1$. This result implies the following proposition.

Proposition 2.1. *For $\bar{\theta} - \underline{\theta} > 0$ and $\rho > 0$, there is always a threshold value $m^* \in (0, \infty)$ for*

[17] See Appendix Section A.1.2 for the proof.

inspection costs such that average procurement costs in both systems are the same. This point is the cut-off at which the buyer switches systems: the American system is chosen for $m < m^*$, and the "Japanese" system is chosen for $m > m^*$.

Proof. See Appendix A.1.3. □

This proposition highlights that changes in the arrival rate of trade wars may endogenously affect the choice of procurement system. Starting at a level of inspection cost m slightly below m^*, a reduction in ρ lowers average costs under the "Japanese" system – while also lowering m^* – and may cause the buyer to switch from the "American" to the "Japanese" system. This setting contrasts with existing studies of relational contracts in trade, where exogenous heterogeneity in discount rates determine relationship-based transactions (Kamal and Tang, 2015; Defever, Fischer, and Suedekum, 2016; Kukharskyy, 2016).[18] In our framework, buyers endogenously determine the effective discount rate of rx_x/q by choosing the optimal procurement system and order size in response to inspection costs and the probability of a future trade conflict.

To map the choice of procurement system into observable trade flows, we examine how order size, frequency, and unit values differ across the two systems as a function of the trade war arrival rate ρ and the inspection cost m for a given q. We restrict our attention to a setting where buyers make a purchase at least once per period, $x^* \leq q$, and where discount rates are small, i.e., $0 < \frac{rx_s}{q} < 1$.

Proposition 2.2. *An increase in the probability of a trade war, which increases ρ, raises the unit value per shipment and reduces the size of shipments (i.e., raises shipment frequency) in the "Japanese" system. An increase in the inspection cost m lowers the unit value per shipment and raises the size of shipments (i.e, reduces shipment frequency) in the "American" system.*

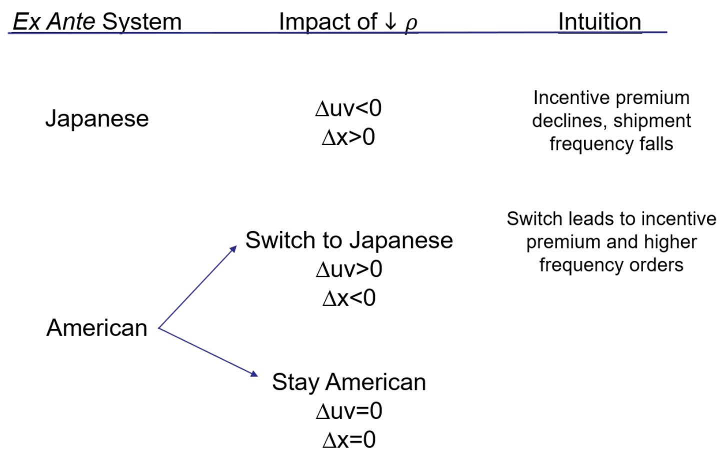
Proof. See Appendix A.1.4. □

Under the "Japanese" procurement system an increase in the separation rate ρ causes sellers to demand a greater premium to maintain quality. As a result, variable procurement costs increase and buyers re-optimize by lowering shipment sizes (i.e., raising shipping frequency). Given that fixed per-shipment costs are spread over smaller shipment sizes, the increase in ρ causes unit values to increase. Procurement patterns under the "American" system are unaffected by an increase in the separation rate. In contrast, an increase in the inspection cost m raises fixed per-shipment costs under the "American" system, and buyers re-optimize by increasing per-shipment quantities (i.e., decreasing shipping frequency). As a corollary, unit values must go down in the "American" system since fixed costs are spread over more units.

We can use these results to rank shipping frequencies and unit values across the two systems. Starting from the case $\bar{\theta} - \underline{\theta} = 0$ and $m_A = 0$, where the "American" and "Japanese" procurement systems are identical, an increase in $\bar{\theta} - \underline{\theta}$ raises variable shipment costs under the "Japanese" system due to the larger incentive premium. Buyers re-optimize by increasing the shipping frequency and

[18] Defever, Fischer, and Suedekum (2016), for example, study a setup in which a buyer and a seller interact repeatedly and choose how much to invest into their relationship. They analyze different values of firms' discount rates and show that a cooperative equilibrium can only exist if both firms are sufficiently patient.

Figure 2: Impact of A Decline in ρ



Notes: Figure illustrates the impact of a change in the arrival rate of a trade war, ρ on shipment prices (p) and quantities (x) under both systems.

lowering the shipment size, which raises unit values since fixed costs are spread over fewer units. Under the "American" system, by Proposition 2.2, an increase in inspection costs leads to a rise in shipment size, and hence shipping frequency and unit values decrease. For the case of $\bar{\theta} - \underline{\theta} > 0$ and $m \geq 0$, it must therefore be true that shipping sizes are greater in the "American" system. Furthermore, unit values are greater in the "Japanese" procurement system compared to "American" procurement system. This reasoning forms the basis of our third proposition.

Proposition 2.3. *Batch sizes in the "American" system are greater than in the "Japanese" system, $x_A^* > x_J^*$, and unit values in the "Japanese" system are greater than in the "American" system, $v_J(x_J, \bar{\theta})/x_J > v_A(x_A, \bar{\theta})/x_A$.*

Proof. See Appendix A.1.5. □

Proposition 2.2 and Proposition 2.3, and Figure 2.2, illustrate that the effect of a decrease in the probability of a future trade conflict, ρ, depends on whether the adjustment takes place within the "Japanese" system or via a switch from the "American" to the "Japanese" system. If adjustment is within the "Japanese" system, then by Proposition 2.2, unit values should fall and shipment size should increase (i.e., shipping frequency should decline) in response to a decline in the probability of a trade war. We would expect this case to be more likely for the U.S. if prior to a decline in the probability of a trade war most U.S. trade relationships were already "Japanese". In contrast, if "American" procurement were prevalent in U.S. trade, then by Proposition 2.3, we would expect a decrease in the probability of a future trade conflict to lead some relationships to switch from "American" to "Japanese" procurement, which would lead to higher unit values and smaller shipment sizes (i.e., greater shipping frequencies).

In the next section, we show that "Japanese" importing from China is relatively rare in particular in the earlier part of the sample period. Consistent with this finding, we show in Section 5 that a policy change which reduced uncertainty with respect to shipments from China led to a change in procurement patterns consistent with a switching of systems rather than adjustment within the "Japanese" system.

3 Data and Description of Procurement Patterns

In the next two sections, we test the implications of our theory using confidential data from the U.S. Census Bureau's Longitudinal Foreign Trade Transaction Database (LFTTD). These data track every U.S. import transaction from 1992 to 2016 and include: the dates the shipment left the exporting country and arrived in the United States; identifiers for the U.S. and foreign firm conducting the trade; the shipment's value and quantity; a ten-digit Harmonized System (HS10) code classifying the product traded; the country of origin of the exporter; and the mode of transport.[19]

We refine the raw data using conventional screens, e.g., by focusing on arm's length transactions and removing all observations that do not include or have invalid an importer identifier, an exporter identifier, an HS code, a value, a quantity, or a valid transaction date. We use the concordance developed by Pierce and Schott (2012) to create time-consistent HS codes so that purchases of goods can be tracked over time, and deflate all values using the quarterly GDP deflator of the Bureau of Economic Analysis. Since shipments of the same product between the same buyer and seller spread over multiple containers are recorded as separate transactions, we aggregate the dataset to the weekly level, so that multiple transactions in the same week appear as a single transaction. Finally, to remove unit value outliers, we follow Hallak and Schott (2011) in dropping observations where the unit value is below the 1st or above the 99th percentile within HS10 by country by mode of transportation by quarter cells. For more detail, see Appendix Section B.

Table 1 provides an overview of the data. From 1992 to 2016, the U.S. imported 5.68 trillion dollars worth of goods at arm's length, the majority of which arrived by water (vessel). These imports span 360 thousand unique US importers and just over 5 million unique foreign exporters. Our analysis below will focus on "buyer quadruples" that group shipments of a ten-digit HS product (h) imported by a U.S. importer (m) from origin country (c) shipped via mode of transportation (z).[20] As indicated in the table, there are almost 3 million such $mhcz$ quadruples between 1992 and 2016. The final row of the table reports the number of "buyer-seller relationships" associated with these bins, i.e., the number of $mxhcz$ quintuples, where x denotes the exporter. There are nearly 22 million of these relationships within the 3 million buyer quadruples, or an average of about 7 sellers per $mhcz$ cell.[21]

[19]The four main modes of transportation are vessel, rail, road, and air. We drop the small fraction of transactions that are transported by other means, e.g., hand-carried by passengers. For further information on the LFTTD, see Bernard, Jensen, and Schott (2009) and for detailed information on the foreign firm identifier, see Kamal and Monarch (2018). Though foreign identifiers in the LFTTD include information about country of origin, we continue to mention this dimension of the data for clarity.

[20]We include mode of transportation in defining these bins to mitigate the influence of spurious sources of variation – e.g., product quality – that might differ across product varieties shipped using different methods.

[21]We realize that referring to "$mhcz$ quadruples" and "$mxhcz$ quintuples" is awkward but have retained this language for its precision. In the data, a given seller (i.e., exporter) may supply a particular HS code to multiple buyers (i.e., importers). To match theory and data, we interpret this behavior as sellers producing different varieties within HS codes for each buyer without any costs to the buyer or seller beyond those described in Section 2.1. Under our contestable market assumption, importers trying to sell a particular final good variety at a price above average cost face competition

Table 1: U.S. Import Transaction Summary Statistics

Total Imports ($\$Bill$)	5,680
Vessel Imports ($\$Bill$)	4,030
Air Imports ($\$Bill$)	988
Unique Importers (m)	360,000
Unique Exporters (x)	5,037,000
Unique Importer-Product-Country-Mode Quadruples ($mhcz$)	2,966,000
Unigue Exporter-Importer-Product-Country-Mode Relationship Quintuples ($mxchz$)	21,700,000

Source: LFTTD and authors' calculations. Table summarizes U.S. arm's-length imports from 1992 to 2016. Import values are in billions of real 2009 dollars. Vessel imports refer to imports arriving over water. Final four rows of tables provide counts of unique importers, exporters, buyer quadruples, i.e., U.S. importer by HS product by origin country by mode of transport cells, and buyer-seller relationships, i.e., U.S. importer by foreign exporter by HS product by origin country by mode of transport cells

Table 2 summarizes the $mhcz$ quadruples, which are the focus of our study in the next section, in greater detail. The first four rows of the table reveal that from 1992 to 2016, the average $mhcz$ bin traded 1.9 million dollars, lasted for 304 weeks and encompassed 39 shipments across 7 sellers. Rows 5 through 7 highlight "procurement patterns," showing that average value per shipment (VPS_{mhcz}), weeks between shipments (WBS_{mhcz}), and buyer-seller relationship length ($length_{mhcz}$) averaged 36 thousand dollars, 24 weeks and 181 weeks, respectively.[22] In each case, the large standard deviations compared to the means indicate that buyer quadruples exhibit substantial heterogeneity. Appendix C provides more details on how all variables are constructed.

Table 2: Attributes of $mhcz$ Quadruplets

	Mean	*Standard Deviation*
Total Value Traded ($)	1,914,000	36,300,000
Length Between Buyer's First and Last Shipment (Weeks)	304.3	266
Total Shipments	38.6	157.9
Number of Sellers (x)	7.3	25.5
Value per Shipment (VPS), ($)	35,910	386,100
Weeks Between Shipments (WBS)	23.5	28.5
Average Relationship Length in Weeks ($length$)	180.8	154.7
Ratio of Sellers to Shipments (SPS)	0.334	0.241

Source: LFTTD and authors' calculations. Table reports the mean and standard deviation across importer (m) by country (c) by ten-digit Harmonized System category (h) by mode of transport (z) quadruplets during our 1992 to 2016 sample period. Observations are restricted to quadruplets with more than one transaction.

from other importers able to offer the same variety using an input from either the same or a different foreign seller. We assume that "American" buyers can procure their variety from different sellers over time, and that different buyers procuring the same product from the same seller might use different procurement systems because inspection costs can vary by variety within a product.

[22]While below we also analyze quantity per shipment (QPS_{mhcz}) and unit value per shipment (UV_{mhcz}), we do not summarize these here due to differences in the unit of quantity across products reported in U.S. import data. We note that the relationship lengths can be subject to both left and right censoring around the beginning and end of our period of analysis.

As discussed in Section 2, a key characteristic of "Japanese" sourcing in the model is the presence of long-term relationships between a buyer and seller, in which the buyer incentivizes the seller to provide goods of a sufficient quality by paying a premium over the seller's costs. As a result, buyers engaged in "Japanese" sourcing are expected to have fewer sellers for a particular good than those sourcing under the "American" system.[23] We build on this insight and take the number of sellers used at the level of the importer-product-country-mode ($mhcz$) quadruple as an observable measure of the extent of "Japanese" sourcing. We normalize this measure by the number of shipments to account for the fact that some buyers receive many shipments, and have many opportunities to source from a different seller, while others receive far fewer. Hence, we compute sellers per shipment (SPS_{mhcz}) as

$$SPS_{mhcz} = \frac{Sellers_{mhcz}}{Shipments_{mhcz}}. \tag{7}$$

Lower SPS_{mhcz} indicates more "Japanese" sourcing, while higher ratios suggest more "American" sourcing.

The final row of Table 2 reveals that the mean ratio of sellers to shipments across buyer quadruples is 0.33, with a standard deviation of 0.24. The kernel density reported in Figure 3 provides a more complete description of the distribution of this attribute. As indicated in the figure, most buyer quadruples have a relatively small ratio of sellers to shipments. However, observations in the right tail approach a value of 1, i.e., a different seller for each shipment.[24]

We provide statistics on SPS_{mhcz} by source country in Table 3. The first two columns report a value-weighted average of SPS_{mhcz} for two different time periods: 1995-2000 and 2002-2007.[25] We present two different periods to provide intuition for the evolution of SPS_{mhcz} over time, which will be useful in Section 5 below. As indicated in the first column of the table, we find that the average number of suppliers per shipment is lowest for U.S. imports from Mexico and Japan, with the latter perhaps reflecting the influence of large "Japanese" multinationals like Toyota producing in the United States (Boehm, Flaaen, and Pandalai-Nayar, 2020). It is also relatively low for Taiwan and Canada, and relatively high for France, China, Brazil, and Germany, all of which have values above the rest of the world.

Results in the second column reveal that, over time, the average SPS_{mhcz} generally falls, with imports from Japan and the UK being exceptions. The largest decreases exhibited, both in levels and percent growth, are for Mexico, China and Brazil.[26] We return to the large fall in SPS_{mhcz} for China below and examine to what extent it is the result of a reduction in trade policy uncertainty with respect to shipments from China that occurs in 2001.

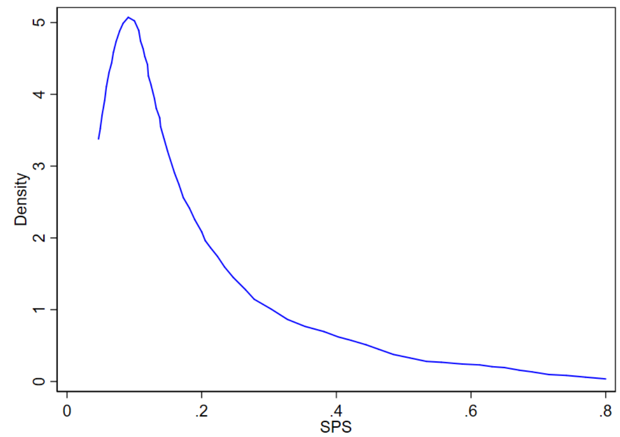
[23]In the model, "Japanese" buyers have one seller for a given product. In practice, "Japanese" buyers might use more than one seller if, for example, they experiment before settling with a long-term partner. Similarly, "American" buyers might be observed to have only one seller if that seller repeatedly offers the lowest price. Nevertheless, we show below that the number of sellers used, normalized by the number of shipments, is correlated with shipment patterns in a manner consistent with the model.

[24]Consistent with the regression analysis below, we consider only $mhcz$ quadruples that have conducted at least 5 transactions in our sample period. Our theory requires that we observe repeated shipments to learn about the procurement system.

[25]Some quadruples extend through both periods, others show up in just one or the other.

[26]The relatively large decline of suppliers per shipment from Mexico may be related to increasingly close supply-chain integration with the United States after NAFTA, including maquiladora trade.

Figure 3: Distribution of Sellers Per Shipment (SPS) Across Relationships, 1992 to 2016



Source: LFTTD and author's calculations. Figure displays the distribution of sellers per shipment (SPS_{mhcz}) across buyer quadruples, i.e., $mhcz$ bins, from 1992 to 2016. The figure was created according to Census Bureau guidelines and omits observations below the 5th percentile and above the 95th percentile.

To examine the *share* of relationships most likely to be "Japanese", in columns 3 and 4 we compute the value-weighted fraction of buyer quadruples whose SPS_{mhcz} falls within the first quartile of the sellers per shipment distribution for their product-mode (hz) bin in the *first* period. The unweighted share of these "Japanese" buyer quadruples within each hz bin in the first period will be 25 percent *worldwide* by construction, but can vary across countries in that period. In the second period, the share of such buyer quadruples worldwide need not be 25 percent. While the cutoff is arbitrary, this variable provide a rough measure of how many quadruples might be "Japanese". We find that, even though the number of "Japanese" quadruples is only 25 percent worldwide, the value of "Japanese" imports accounts for the majority of trade for all countries. For Japan and Mexico, they account for about three quarters of imports. Similar to our findings for SPS_{mhcz}, the share of "Japanese" imports is increasing over time for most countries, with the strongest increases recorded for Brazil, China, and Mexico.

Table 3: "Japanese" Relationships by Country

	Mean SPS		$SPS = Q1$ Import Share	
Country	1995-2000	2002-2007	1995-2000	2002-2007
Mexico	0.093	0.069	0.750	0.869
Japan	0.118	0.131	0.756	0.725
Taiwan	0.138	0.122	0.711	0.743
Canada	0.147	0.130	0.602	0.667
United Kingdom	0.157	0.229	0.717	0.519
South Korea	0.166	0.141	0.656	0.724
Rest of the World	*0.187*	*0.160*	*0.625*	*0.678*
France	0.192	0.168	0.627	0.667
China	0.192	0.150	0.582	0.693
Brazil	0.197	0.158	0.576	0.706
Germany	0.203	0.179	0.582	0.606

Source: LFTTD and authors' calculations. First two columns report the weighted average sellers per shipment (SPS_{mhcz}) across buyer quadruples by country and period, where import values are used as weights. The second two columns report the share of the value of U.S. imports accounted for by quadruples with SPS_{mhcz} in the first quartile of the distribution of SPS_{mhcz} within product-mode in the first period. Rows of the table are sorted by the first column.

4 American vs "Japanese" Buyers

In this section we show that the purchasing patterns of buyers with higher versus lower ratios of sellers per shipments, SPS_{mhcz}, are consistent with the predictions of our model. We demonstrate that U.S. importers with lower SPS_{mhcz} – i.e., those more likely to be "Japanese" – do indeed order smaller quantities per shipment, order more frequently, and pay a larger average price (unit value) per shipment, than more "American" buyers with higher SPS_{mhcz}. These results justify our use of SPS_{mhcz} as a measure of the extent of "Japanese" versus "American" sourcing in our analysis of PNTR in the next section.

4.1 A vs J Among Buyer Quadruples

We first evaluate the appropriateness of using SPS_{mhcz} as a measure of the extent of "Japanese" sourcing by running an $mhcz$-level OLS regression,

$$ln(\overline{Y}_{mhcz}) = \beta_1 ln(SPS_{mhcz}) + \beta_2 \ln(QPW_{mhcz}) + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{hcz} + \epsilon_{mhcz}. \quad (8)$$

The dependent variable, $\overline{Y}_{mhcz}$, represents averages of procurement patterns including quantity per shipment (QPS_{mhcz}), weeks between shipments (WBS_{mhcz}), and price per unit (UV_{mhcz}) across all transactions of a given $mhcz$ quadruple in our dataset. In line with our assumption holding quantity fixed in Section 2, we condition on buyers' total order "flow" by controlling for the total quantity imported by the buyer quadruple over its entire lifetime divided by its overall length, in weeks, QPW_{mhcz}.[27] We control for quadruples' first and last weeks of trade, beg_{mhcz} and end_{mhcz}, to capture potential effects of trading in a specific time period—such as a particular stage in the business cycle—and duration effects.[28] Our regression also includes product by country by mode of transportation fixed effects (λ_{hcz}). These fixed effects capture time-invariant characteristics of trade along these dimensions such as distance, transit time, or level of transportation infrastructure. We describe the construction of all variables in detail in Appendix C. The sample period is 1992 to 2016, and standard errors are clustered at the country level. We exclude buyer quadruples with fewer than 5 shipments since our theory requires that we observe repeated shipments to learn about the procurement system. Moreover, quadruples with fewer shipments might represent importers trying out a new product or other idiosyncrasies.[29]

Results for specification (8) are reported in Table 4. Consistent with Proposition 2.3, we find that quadruples with higher SPS_{mhcz}, i.e., those that are more "American", receive shipments that are larger, less frequent, and lower in price. The coefficient estimates indicate that a one standard deviation increase in sellers per shipment (0.24) is associated with a 0.10 log point rise in quantity per shipment, a 0.11 log point increase in weeks between shipment, and a 0.03 log point decline in price.[30]

[27]We normalize the total quantity traded by the number of weeks since it is straightforward to implement in our weekly dataset. An alternative would be to use the annual quantity traded, which is simply $52 \cdot QPW_{mhcz}$.

[28]beg_{mhcz} and end_{mhcz} are continuous variables indicating the week numbers that the relationship commences and ceases.

[29]In Appendix D.1, we show that results are qualitatively identical for a cutoff of 10 shipments.

[30]As discussed later, in Section 6.2, the coefficients on QPW_{mhcz} are consistent with Proposition 6.1. In both procurement systems, an increase in the total procured quantity increases shipment size, lowers the number of weeks between shipments and hence raises shipment frequency, and lowers unit import values.

Table 4: A vs J Classification Regression Across $mhcz$ Quadruples

Dep. var.	(1)	(2)	(3)
	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$
$\log(SPS_{mhcz})$	0.418^{***}	0.452^{***}	-0.123^{***}
	0.014	0.014	0.019
$\log(QPW_{mhcz})$	0.701^{***}	-0.308^{***}	-0.287^{***}
	0.010	0.010	0.009
Observations	2,966,000	2,966,000	2,966,000
Fixed effects	hcz	hcz	hcz
R-squared	0.947	0.674	0.845
Controls	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). (QPS_{mhcz}), (WBS_{mhcz}), and (P_{mhcz}) are average quantity per shipment, average weeks between shipment, and average unit value. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country (c) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

The computation of SPS_{mhcz} for each $mhcz$ quadruple is motivated by the assumption that buyers can optimize and choose different procurement systems within each hcz bin. However, in principle, buyers could optimize across modes of transportation, and choose one procurement system per country-product, i.e., across modes. They could also optimize across countries, and choose only one system per mh. As a result, SPS might alternatively be computed at these more aggregate levels. In Appendix D.2, we therefore re-run specification (8) with SPS computed at the level of buyer triples (SPS_{mhc}) and buyer doubles (SPS_{mh}). Results are qualitatively similar to the baseline results at the buyer quadruple level (SPS_{mhcz}).

In Table 5, we consider a related specification that relaxes the restriction of a linear relationship between shipment characteristics and SPS_{mhcz} by estimating

$$\overline{Y}_{mhcz} = \sum_{n\in\{2,3,4\}} \alpha_n 1\{SPS_{mhcz} = Q_n\} + \beta_2 \ln(QPW_{mhcz}) + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{hcz} + \epsilon_{mhcz}, \tag{9}$$

where Q_n index quartiles of the SPS_{mhcz} distribution and $1\{SPS_{mhcz} = Q_n\}$ are dummies that are equal to one if SPS_{mhcz} is in quartile n. We compute the quartile cutoffs by looking across SPS_{mhcz} within hcz bins, and assign each buyer quadruple to its respective quartile. The first quartile – consisting of the most "Japanese" quadruples – is the excluded category. $\overline{Y}_{mhcz}$ continues to represent average procurement patterns, and the remaining controls are the same as in equation

(8). As before, the sample period is 1992 to 2016, standard errors are clustered at the c level, and observations are restricted to $mhcz$ bins with at least 5 transactions.

Results continue to justify use of SPS_{mhcz} as an indicator of procurement system. As illustrated in the first and second columns of Table 5, we find that the SPS_{mhcz} point estimates are positive and statistically significant at conventional levels for both quantity per shipment and weeks between shipments for all quartiles.[31] Moreover, the coefficients rise monotonically from quartile 1 to quartile 4, and are statistically different from one another, consistent with the quadruples in those quartiles being increasingly "American." Coefficient estimates indicate that average quantity per shipment and weeks per shipment for U.S. importers in the fourth quartile are roughly 0.79 and 0.86 log points higher than for those in the first quartile. Coefficient estimates in column 3 of Table 5 provide similar support for the model, indicating that the average price declines with seller per shipment quartile. For example, the average price for the fourth quartile of the SPS_{mhcz} distribution is roughly 0.23 log points lower than that for the first quartile.[32]

Table 5: A vs J Quartile Classification Regression Across $mhcz$ Quadruples

Dep. var.	(1)	(2)	(3)
	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$
$(SPS_{mhcz} = Q2)$	0.328^{***}	0.350^{***}	-0.117^{***}
	0.003	0.003	0.004
$(SPS_{mhcz} = Q3)$	0.552^{***}	0.591^{***}	-0.179^{***}
	0.004	0.004	0.005
$(SPS_{mhcz} = Q4)$	0.792^{***}	0.856^{***}	-0.226^{***}
	0.006	0.006	0.007
$\log(QPW_{mhcz})$	0.687^{***}	-0.323^{***}	-0.282^{***}
	0.003	0.003	0.005
Observations	$2,966,000$	$2,966,000$	$2,966,000$
R-squared	0.945	0.661	0.845
Fixed effects	hcz	hcz	hcz
Controls	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). (QPS_{mhcz}), (WBS_{mhcz}),and (P_{mhcz}) are average quantity per shipment, average weeks between shipment, and average unit value. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country (c) bin are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

[31] As above, and described in section 6.2, the coefficients on QPW_{mhcz} are also consistent with Proposition 6.1.

[32] In Appendix Section D.3, we analyze procurement patterns separately for shipments by air and by vessel, and find that SPS_{mhcz} classifies procurement systems within each of these modes of transportation. In Appendix Section D.4, we analyze procurement patterns within $mxhcz$ buyer-seller relationships, and again find that SPS_{mhcz} is related to procurement patterns in the manner predicted by the model.

In the next subsection, we present additional evidence of "Japanese" procurement by considering relationship length as an additional indicator of "Japanese" procurement.[33]

4.2 Relationship Length

Buyers under the "Japanese" system rely on repeat purchases from the same seller, while buyers under the "American" system choose the (potentially different) lowest cost supplier for each transaction. An implication of our theory is therefore that buyers using "Japanese" procurement should have, on average, longer relationships with their suppliers. To investigate this link, we construct a new variable, average relationship length ($length_{mhcz}$), which captures the average age of the $mxhcz$ buyer-seller relationships within buyer quadruples. We compute this average in two steps. First, for each $mxhcz$ transaction, we compute the number of weeks passed since the first transaction of any good using any mode between the buyer m and seller x. For each $mhcz$ buyer quadruple, we then take the average of these lengths across the $mxhcz$ transactions within it.[34] This average allows for the possibility that buyers already sourcing one product from a given supplier, or already using a different mode of transportation with that seller, may structure relational contracts to encompass all products.[35]

We estimate the same regressions as in the previous subsections but now use $length_{mhcz}$ as the dependent variable. The results, reported in Table 6, show that buyer quadruples with lower ratios of suppliers per shipments tend to have longer relationships. The estimate in Column 1, from specification (8), shows that a one standard deviation decrease in the ratio of sellers per shipments is associated with a 0.14 log point increase in average relationship length. Column 2, based on specification (9), illustrates that the average relationship length for U.S. importers in the fourth quartile of SPS_{mhcz} is roughly 1.20 log points lower than that in the first quartile.

The preceding results demonstrate that key implications of the model are confirmed in the data and that the ratio of sellers per shipment is an appropriate measure for the extent of Japanese-style sourcing. In particular, we showed that buyers with lower SPS_{mhcz} exhibit lower quantity per shipment and weeks between shipments, as well as higher average prices, consistent with Proposition 2.3. We also showed that a lower SPS_{mhcz} is associated with longer relationships, as implied by the theory.

[33] In Appendix Section D.5, we show that buyers of differentiated products tend to puchase smaller quantities more frequently and at a higher price. To the extent that differentiated products have larger inspection costs, these relationships are consistent with Propositions 2.1 and Proposition 2.3.

[34] Using this approach, $length_{mhcz}$ is zero if an mx buyer-seller pair trades only a single good once, and increases if that good is traded more often.

[35] We note that we find similar results – available on request – if we instead define relationship length for $mxhcz$ quintuples. Further details on the computation of relationship length are in contained in Appendix Section C.

Table 6: A vs J Classification Regressions, Using Relationship Length

Dep. var.	(1)	(2)
	$\log(length_{mhcz})$	$\log(length_{mhcz})$
$\log(SPS_{mhcz})$	-0.594^{***}	
	0.013	
$(SPS_{mhcz} = Q2)$		-0.369^{***}
		0.002
$(SPS_{mhcz} = Q3)$		-0.679^{***}
		0.003
$(SPS_{mhcz} = Q4)$		-1.204^{***}
		0.005
$\log(QPW_{mhcz})$	-0.117^{***}	-0.101^{***}
	0.003	0.001
Observations	$2,966,000$	$2,966,000$
R-squared	0.451	0.439
Fixed effects	hcz	hcz
Controls	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing the average relationship length ($length_{mhcz}$) of US importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). The regressions include product by country by mode of transport (hcz) fixed effects. All regressions control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country (c) bin are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

5 The Effect of PNTR on the Choice of Procurement System

A key insight from the model presented in Section 2 is that trade policy can affect buyers' choice of procurement system by affecting the probability of trade wars. In particular, the model suggests that a decrease in the possibility that a trade war will sever relationships can induce buyers and sellers to shift away from "American" to "Japanese" procurement. The inverse result is also true, making the model's implications especially pertinent in the current environment, in which views toward trade and policy changes have raised the risk of trade wars.

In this section, we examine the model's implications using a plausibly exogenous change in U.S. trade policy: the U.S. granting of permanent normal trade relations (PNTR) to China, which substantially reduced the possibility of a trade war between the two countries. We first provide some background on the policy before connecting the policy change to a shift towards "Japanese" sourcing.

5.1 Description of PNTR

U.S. imports from non-market economies such as China are generally subject to relatively high "column two" tariff rates originally set under the Smoot-Hawley Tariff Act of 1930, as opposed to the generally low Normal Trade Relations (NTR) tariff rates the U.S. offers to trading partners that are members of the World Trade Organization (WTO). A provision of U.S. trade law, however, allows imports from non-market economies to enter the United States under NTR tariffs subject to *annual* approval by both the President and Congress. Chinese imports first began entering the United States under this provision in 1980, after the warming of bilateral relations.

Annual approval became controversial and less certain after the Tiananmen Square incident in 1989, however, and this uncertainty continued throughout the 1990s. During this time, firms engaged in or considering U.S.-China trade faced the possibility, each year, of substantial tariff increases if China's NTR status was not re-approved. The magnitude of these potential tariff increases—32 percentage points for the average product—make clear that some buyer-seller relationships that were profitable under NTR tariff rates would not be profitable under a shift to "column two" tariffs. Indeed, Pierce and Schott (2016) document extensive discussion by U.S. firms of the trade-dampening effects of this uncertainty in the 1990s, and Alessandria, Khan, and Khederlarian (2019) show that it reduced U.S. imports from China, while also driving intra-year seasonal patterns in imports.

When the U.S. granted PNTR to China in 2001, it locked in NTR rates, eliminating the need for annual renewals and the potential for relationship-severing tariff increases. Pierce and Schott (2016) show that this elimination of uncertainty was associated with reductions in U.S. manufacturing employment as imports from China increased, and Handley and Limão (2017) estimate that the elimination of uncertainty was equivalent to a 13 percentage point decline in tariff rates. Note that this policy change did not affect actual tariffs, and would therefore have no effect on trade flows in models such as Eaton and Kortum (2002), while in our framework it affects ρ and hence procurement costs under the "Japanese" system.

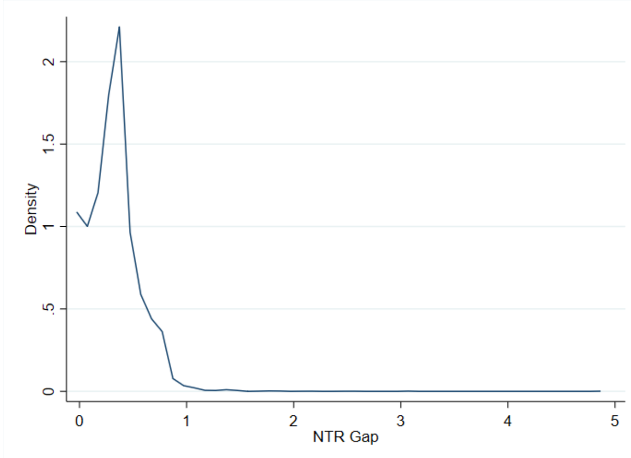
This policy environment provides a useful opportunity for testing Proposition 2.1, which states that a decrease in the probability of a trade war leads to the adoption of more "Japanese" sourcing. We follow Pierce and Schott (2016) in defining a products' exposure to PNTR as the difference between the non-NTR rate to which its tariff could have risen before PNTR and the lower NTR rate that was locked in by the policy change,

$$NTR\,Gap_h = Non\,NTR\,Rate_h - NTR\,Rate_h. \tag{10}$$

We compute these gaps as of 1999, the year before the change in policy, using *ad valorem* equivalent tariff rates provided by Feenstra, Romalis, and Schott (2002). As indicated in Figure 4, these gaps vary widely across products, and have a mean and standard deviation of 0.32 and 0.23, respectively.

Our identification strategy exploits this variation in the NTR gap to determine whether U.S.-China procurement patterns change relative to procurement patterns with exporters from other source countries (first difference) after the change in U.S. policy is implemented (second difference) for products with higher NTR gaps (third difference). The last difference captures the fact that products

Figure 4: Distribution of the *NTR Gap*



Source: Feenstra, Romalis, and Schott (2002) and author's calculations. Figure displays the distribution of the $NTR\ Gap_h$, the difference between the relatively low NTR tariff rate that was locked in by PNTR and the higher rate to which U.S. tariffs on Chinese goods might have risen absent the change in policy.

with larger NTR gaps experience a larger decline in the relationship termination probability than products with smaller gaps. As described above, this relationship termination probability depends on the change in the probability of China's NTR status – which is identical for all products – and the increase in tariff rates that could have occurred before PNTR, which varies by product. We expect the largest shifts toward "Japanese"-style procurement after PNTR to occur in U.S. imports of high-NTR gap products from China.

5.2 Response to PNTR Among Continuing $mxhcz$ Quintuples

We start by analyzing procurement patterns in a balanced panel of *continuing* buyer-seller $mxhcz$ quintuples, i.e., within those buyer-seller relationships that trade both before and after PNTR. Our theory assumes a stable environment of repeated buyer-seller interactions and does not explicitly model the entry and exit of suppliers – therefore, this balanced panel is a natural starting point to test our model's predictions. We subsequently expand the analysis to study new buyer-seller relationships.[36]

Our specification examines the relationship between PNTR and procurement characteristics across two five-year periods (p) straddling the implementation of PNTR in 2001, 1995 to 2000 and 2002 to 2007[37]:

[36] Pierce and Schott (2016); Amiti, Dai, Feenstra, and Romalis (2020) show that PNTR led to a significant increase in Chinese firms exporting to the United States, and hence this margin could affect the aggregate effects.

[37] We choose these periods to have symmetric length around the policy change, which took effect at the end of 2001, and so that the latter window ends before the onset of the Great Recession. In Appendix Section E.1, we show that we find substantially similar results using different time periods.

$$ln(Y_{mxhczp}) = \beta_1 1\{p = Post\} * 1\{c = China\} * NTR\,Gap_h + \beta_2 ln(QPW_{mxhczp}) + \beta_3 \chi_{mxhczp} + \lambda_{mxhcz} + \lambda_p + \epsilon_{mxhczp}. \quad (11)$$

In this $mxhczp$-level OLS regression, Y_{mxhczp} represents buyer-seller relationship-level procurement patterns computed separately for each period – i.e., $\ln(QPS_{mxhczp})$, $\ln(WBS_{mxhczp})$, and $\ln(UV_{mxhczp})$.[38] The triple difference-in-differences (DID) term of interest is an interaction of a dummy for the post period ($1\{p = Post\}$), a dummy for imports from China ($1\{c = China\}$), and the $NTR\ Gap_h$. The variable χ_{mxhczp} represents the full set of interactions of those variables required to identify β_1. The remaining terms on the right-hand side control for the average quantity traded per week in each of the two periods (QPW_{mxhczp}) as well as buyer-seller relationship (λ_{mxhcz}) and period (λ_p) fixed effects. Standard errors are clustered at the country level.

Results, reported in Table 7, indicate that higher exposure to PNTR is associated with a shift toward Japanese-style procurement within existing buyer-seller quintuples. Coefficient estimates in the first two columns show that a one standard deviation increase in the $NTR\ Gap$ induces a relative decline in quantity per shipment and weeks between shipments of 4.5 percent. Moreover, results in column 3 reveal that a one standard deviation increase in exposure to PNTR is associated with a relative increase in price of 2.1 percent. In each case, the findings in Table 7 are consistent with the predictions of Propositions 2.1 and 2.3, indicating a switch from "American" to "Japanese" procurement, as opposed to an adjustment within the "Japanese" system.[39] We provide further evidence for this switch towards the "Japanese" system in the following sections.

One concern with the results in Table 7 is that PNTR may affect the overall quantity traded, which in turn influences the choice of procurement system. We show in Appendix E.3 that our results also hold when we do not include QPW_{mxhcz} as a covariate. Hence, our findings are robust to the endogenous adjustment of total trade in response to PNTR.

[38] Appendix C provides more details on the variables.

[39] Consistent with Proposition 6.1, the coefficient estimates for $ln(QPW_{mxhczp})$ indicate that an increase in the procurement quantity increases the size of shipments, raises shipping frequency, and reduces unit values.

Table 7: Baseline Within *mxhcz* Quintuple PNTR DID Regression

Dep. var.	(1)	(2)	(3)
	$\ln(QPS_{mxhczp})$	$\ln(WBS_{mxhczp})$	$\ln(UV_{mxhczp})$
$Post_p * China_c * NTRGap_h$	-0.1970***	-0.1970***	0.0922***
	0.0156	0.0156	0.0167
$ln(QPW_{mxhczp})$	0.3680***	-0.6320***	-0.1238***
	0.0077	0.0077	0.0113
Observations	439,000	439,000	439,000
R-squared	0.98	0.88	0.98
Fixed effects	$mxhcz, p$	$mxhcz, p$	$mxhcz, p$
Controls	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by exporter by product by country by mode of transport (*mxhcz*) bins on the difference-in-differences term of interest and quantity shipped per week. Pre-and post periods are 1995 to 2000 and 2002 to 2007. QPS_{mxhczp}, WBS_{mxhczp}, and UV_{mxhczp} are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period *p*. All regressions include *mxhcz* and period *p* fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

5.3 Response to PNTR Among New *mxhcz* Quintuples

In this section, we shift to examining the purchasing patterns of *new mxhcz* quintuples formed in the post-PNTR period to quintuples that were new in the pre-period, again using equation (11). While our theory requires repeated interactions between buyers and sellers, it does not require the relationships to be long-established. PNTR should therefore affect procurement patterns for new relationships in a similar way as for established relationships.

In the pre-PNTR period, we define a new relationship as an *mx* pair that appears for the first time in 1995 to 2000, i.e., it is not present in any prior years for which data are available (from 1992 to 1994). Likewise, a new relationship in the 2002 to 2007 period is an *mx* pair not observed previously, from 1992 to 2001. As above, the estimation is performed at the *mxchzp* level, and standard errors are clustered at the country level. Here, however, because we are comparing different buyer-seller relationships in each period, we cannot include *mxhcz* fixed effects, and instead include buyer (*mhcz*), exporter (x), and period (p) fixed effects. As a consequence, our regression focuses on the new relationships formed by buyers and sellers that exist in both time periods, but who form new relationships across time periods.

Results, reported in Table 8, are also consistent with the model's predictions. As indicated in the table, we find that buyer-seller relationships formed after PNTR trading goods with greater exposure to the policy change exhibit relatively smaller and more frequent shipments, at relatively higher prices. The coefficient estimates reported in the table indicate that a one standard deviation increase in exposure is associated with a 2.7 percent decline in shipment size and frequency, and a 2.1 percent rise in price.[40]

[40] In Appendix E.4, we re-run specification (11) using all relationships (i.e. not only new relationships) that appear

Table 8: New *mxhcz* Quintuple PNTR DID Regression

Dep. var.	(1)	(2)	(3)
	$\ln(QPS_{mxhczp})$	$\ln(WBS_{mxhcz})$	$\ln(UV_{mxhcz})$
$Post_p * China_c * NTRGap_p$	-0.1164***	-0.1164***	0.0895***
	0.0247	0.0247	0.0316
$ln(QPW_{mxhcz})$	0.4087***	-0.5913***	-0.1291***
	0.0117	0.0117	0.0163
Observations	3,184,000	3,184,000	3,184,000
R-squared	0.9662	0.8337	0.9715
Fixed effects	$mhcz, x, p$	$mhcz, x, p$	$mhcz, x, p$
Controls	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of comparing new buyer-seller relationships in the pre-versus post-PNTR period. Pre-and post periods are 1995 to 2000 and 2002 to 2007. New relationships are defined as mx pairs appear for the first time in each period. (QPS_{mxhczp}), (WBS_{mxhczp}), and (UV_{mxhczp}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period p. All regressions include $mhcz$, x and period p fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

5.4 Response to PNTR in Supplier-Shipment Ratio and "Japanese" Indicator

We provide further evidence for the shift toward "Japanese" procurement by examining the relationship between PNTR and the two measures of Japanese procurement described above, the ratio of sellers per shipments (SPS) and the indicator for Japanese procurement (J).[41] These attributes cannot be examined in the $mxhcz$-level analyses in the previous two sections, because they do not vary within buyer-seller pairs. Therefore, we consider the relationship between PNTR and these attributes at two higher levels of aggregation, the $mhcz$ quadruple level, and the hcz triple level. We aggregate both SPS and J to the hcz level by taking their weighted average across the quadruples contained in the hcz cell, using import weights.

Our estimating equation is a variant of equation (11), with the fixed effects λ adjusted to the relevant level of aggregation. Note that identification in the $mhcz$ level analysis includes the impact of continuing firms and new exporters, while identification in the hcz level analysis includes the impact of continuing firms, new exporters, *and* new importers.

Results, reported in Table 9 indicate a shift toward Japanese-style sourcing that becomes more precisely estimated as higher levels of aggregation—and therefore additional margins of firm entry—are considered. As shown in columns 1 and 2 of the table, analysis at the $mhcz$ level indicates that buyer quadruples more exposed to the granting of PNTR to China exhibit lower, but statistically insignificant, ratios of sellers per shipment (SPS_{mhzp}) and a higher probability of Japanese procurement (J_{mhzp}), relative to less-exposed buyers. At the hcz level, where entry of new importers is

in either the pre-PNTR or the post-PNTR period, provided that the buyer quadruple and the seller appear in both, by using separate $mhcz$ and x fixed effects. The results are similar.

[41] Recall that this "Japanese" indicator is equal to one if SPS_{mhcz} falls into the first quartile of its distribution within the associated hz bin, where the distribution is held fixed from the pre-PNTR period.

also considered, higher exposure to PNTR is associated with both a lower ratio of sellers per shipment (column 3) and higher probability of Japanese procurement (column 4), with each relationship substantially more precisely estimated than at the more disaggregated level.[42]

Combined with Pierce and Schott (2016)'s finding that PNTR is associated with increases in importer-exporter pairs, these results highlight the relevance of new exporters and importers in the post-PNTR shift toward Japanese procurement. In other words, beyond lowering procurement costs for firms already engaged in US-China before PNTR, the policy change may have facilitated the entry of a new set of firms that were newly able to sustain the long-term relationships that characterize the "Japanese" system.

Table 9: Within-Importer PNTR Regression, Buyer Characteristics

Dep. var.	(1)	(2)	(3)	(4)
	$\ln(SPS_{mhcz})$	J_{mhcz}	$\ln(SPS_{mhcz})$	J_{mhcz}
$Post_p * China_c * NTRGap_p$	-0.0064	0.0414*	-0.0205**	0.0582***
	0.0309	0.0218	0.0090	0.0189
$ln(QPW_{mhcz})$	0.4361***	-0.5639***	-0.0623***	-0.0625***
	0.0172	0.0172	0.0020	0.0021
Observations	738,000	291,000	368,000	28,500
R-squared	0.7727	0.6752	0.6949	0.5308
Fixed effects	$mhcz, p$	$mhcz, p$	hcz, p	hcz, p
Controls	Yes	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by product by country by mode of transport ($mhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre-PNTR period is 1995 to 2000. First and second two columns summarize results when the post period is 2002 to 2007, and 2004 to 2009, respectively. All regressions include $mhcz$ and period p fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Overall, the results in this section provide evidence consistent with one of the key implications of the model, namely that a change in trade policy that changes the probability of relationship continuation can induce changes in firms' sourcing behavior. Across existing and new relationships and defined at different levels of aggregation, we consistently find that the granting of PNTR to China in 2001, which reduced the probability of tariff increases, was associated with a switch toward "Japanese" sourcing. In addition to providing empirical support for the model's framework, the results also suggest important implications of more recent U.S. trade policy. In particular, the results suggest that the more uncertain trade policy environment since the imposition of sizable tariffs by the U.S. and China in 2018 and 2019 is likely already suppressing the formation of long-term "Japanese" relationships.

[42] To analyze the influence of initial buyer experimentation during the years immediately after PNTR on our results, we also consider, in Appendix E.1, similarly constructed outcomes but for a slightly later — 2004 to 2009 — post-PNTR time period. Coefficient estimates for this alternate post period have the same sign patterns, but are larger in absolute magnitude and are more precisely estimated, suggesting adjustment to PNTR may have occurred gradually.

6 Multi-Country Framework with Endogenous Demand

In this section we embed our model of procurement in a general equilibrium theory of trade to quantify how trade patterns and aggregate welfare are affected by access to the two different procurement systems. We start with the multi-country framework of Eaton and Kortum (2002), generalizing their setup to include the possibility of a trade war and downward-sloping average costs.

6.1 Environment

There are N countries, which we index by i and n. Each country is populated by L_n consumers, who purchase a continuous flow of goods from a manufacturing sector and a non-manufacturing sector to maximize a Cobb-Douglas utility of the form $Q_n^{\alpha} Z_n^{1-\alpha}$, where Q_n is the quantity of a composite manufactured good and Z_n is the quantity of a homogeneous, non-manufactured good. The manufactured good is a CES aggregate of a continuum of differentiated products indexed by $\omega \in [0,1]$,

$$Q_n = \left(\int_0^1 q_n(\omega)^{(\sigma-1)/\sigma} d\omega \right)^{\sigma/(\sigma-1)}, \tag{12}$$

where $\sigma > 0$ is the elasticity of substitution and $q_n(\omega)$ is quantity. Consumers purchase each manufactured good ω from homogeneous buyer firms in their country, which offer the goods at prices $p_n(\omega)$. Country n's price index for manufactured goods is thus

$$p_n = \left(\int p_n(\omega)^{1-\sigma} d\omega \right)^{1/(1-\sigma)}. \tag{13}$$

We assume that each consumer supplies one unit of labor.

Manufactured good ω can be produced by homogeneous seller firms in country i with the linear production function $q = \frac{\Upsilon}{\theta} l$, where l is the quantity of labor used by the seller and $\theta \in \{\underline{\theta}, \bar{\theta}\}$ is quality, as described in Section 2. Labor is paid the wage rate w_n in country n. The productivity $\Upsilon_i(\omega)$ is specific to each origin country-product combination and, as in Eaton and Kortum (2002), drawn from a Frechet distribution according to

$$F_n(\Upsilon) = e^{-T_n \Upsilon^{-\zeta}}, \tag{14}$$

where the country-specific parameter T_n scales the mean of the distribution and ζ scales the variation. The productivity draws are independent across products within each country. Sellers in country n also incur fixed logistic and transportation costs f_n in units of seller country labor for each destination that is supplied. We assume that a country's firms are owned by the household.

Transactions between buyer firms and seller firms take place as described in Section 2. Buyer firms choose whether to purchase goods using "American" or "Japanese" procurement. Buyers using the "American" system need to use an additional $m(\omega)$ labor units, drawn from a distribution $G(m)$, to inspect the quality of product ω. Buyers choosing the "Japanese" system pay an incentive premium to ensure quality. Sellers obtain profits per order cycle that are equal to the incentive premium under the "Japanese" system and zero under the "American" system. We denote by $\pi^s_{ni,s}(\omega)$ the continuous flow of profits to country i from sales to country n of variety ω under system s.

Non-manufacturing output in country n is produced according to $Z_n = a_n L_n^{NM}$, where a_n is country n's productivity in non-manufacturing and L_n^{NM} is the labor used by the non-manufacturing sector. The non-manufacturing good can be costlessly traded across countries, and is therefore used as numeraire. As in Eaton and Kortum (2002) and Antras, Fort, and Tintelnot (2017), we assume that labor is perfectly mobile between manufacturing and non-manufacturing, and that the non-manufacturing sector is sufficiently large so that at least some of its output is produced.

6.2 Endogenous q in a Contestable Market

Our extended model moves beyond the environment introduced in Section 2 by assuming that total quantity ordered, q, is endogenous. We now construct the equilibrium in the general model, proceeding in two steps. First, in this section, we describe the equilibrium in the single product-destination country market of Section 2 with endogenous q, assuming that the buyer has already chosen the source country and procurement system. In the next section, we then analyze how the seller's country and the procurement system are chosen when we embed the product market in the overall general equilibrium of the economy.[43]

Our first step, Proposition 6.1, shows that batch size and shipping frequency increase with quantity ordered, q. We next show that, as a result, average cost curves in our model are downward sloping in q. Finally, we construct the market equilibrium in the presence of downward sloping average costs.

Proposition 6.1. *An increase in the procurement target q raises batch sizes x_s^* and the shipping frequency q/x_s^* in both systems, and, as a corollary, lowers unit values in both systems.*

Proof. See Appendix A.1.6. □

Intuitively, for a given fixed shipping frequency, firms must increase the batch size x in both systems to meet an increase in q. But by the first-order condition, equation (6), we know that firms trade-off variable procurement costs against fixed per-shipment costs. Therefore, as variable procurement costs increase, buyers respond by spreading the larger quantities over more shipments, raising shipping frequency. As a result, larger quantities purchased lead to both greater shipment sizes, x, but also greater order frequencies. It follows that unit values in both procurement systems decrease, since fixed per-shipment costs are spread over greater per-shipment quantities. Additionally, in the "Japanese" system, an increase in the shipping frequency implies a lower premium to motivate desired quality, which lowers the unit value further.

The comparative statics with respect to q are supported by the empirical estimates in Sections 4 and 5. As indicated in Tables 4, 5, and 7 we find that shipment size is positively related to the quantity shipped per week (QPW), and that both shipping frequency and unit values are negatively related to QPW.

We next show that greater shipment sizes x_s^* as q increases imply downward sloping average cost curves:

[43] We omit unnecessary subscripts here since we focus on a single market.

Lemma 6.2. *At the optimal order size x_s^*, both procurement systems provide economies of scale, i.e., $\frac{\partial AC(x_s^*,q)}{\partial q} < 0$. Moreover, the second derivative of the average cost with respect to q is positive, $\frac{\partial^2 AC(x_s^*(q),q)}{\partial q^2} > 0$, and the average cost in both systems reaches a positive and finite limit as $q \to \infty$.*

Proof. See Appendix A.1.7. □

Downward sloping average cost curves are a key departure of our model from standard trade models, which generally assume constant marginal cost. Our model generates a natural monopoly in the consumer market due to the presence of the fixed cost of ordering. We therefore replace Eaton and Kortum (2002)'s assumption of perfect competition with the assumption that buyers compete in a "contestable" market for consumers. Introduced by Baumol, Panzar, and Willig (1982) and Tirole (1988), contestable markets are a natural extension of Bertrand competition when firms' costs exhibit economies of scale. In a contestable market, there exist several homogeneous competitors whose entry is costless. Due to the economies of scale, in equilibrium it must be the case that a single buyer serves the entire consumer market. Lemma 6.2 indicates that average cost curves are convex, and therefore a demand curve that uniquely intersects the single buyer's optimized average cost curve from above determines a unique, sustainable and feasible equilibrium in the product market, q^*. The buyer prices and supplies the final consumers along its average cost curve. Therefore, no firm can undercut the incumbent. If the buyer prices above average costs, then entrants contest the positive profits, pushing the buyer out and taking over the market. If the buyer prices below average costs, he realizes negative profits. Because consumers are willing to pay prices greater than average costs for $q < q^*$, potential entry forces the incumbent to lower prices and increase quantity such that the market clears where supply equals demand. The buyer is not willing to procure a greater quantity because she would incur losses.

In principle, our CES demand system may intersect the downward sloping average cost curve multiple times. For equilibrium to exist, it must be the case that the demand curve cuts the average cost curve from above at the intersection that determines the greatest equilibrium quantity, q^*_{high}. Intuitively, if the demand curve were above the average cost curve for all $q > q^*_{high}$, then consumers would be willing to buy an infinite quantity of the good when buyer firms set prices equal to average costs. Therefore, under appropriate assumptions on the demand system, the market equilibrium is a corollary of Lemma 6.2.

Corollary 6.2.1. *If markets are contestable and demand intersects average costs from above at q^* and remains below average costs as $q^* < q \to \infty$, then a single buyer procures the product from the seller and distributes it on the consumer market using the buyer's cost minimizing procurement system at optimal shipping frequencies.*

6.3 Overall Equilibrium with Endogenous q

We now embed the product market equilibrium into the equilibrium of the overall economy. Equilibrium requires that buyer firms minimize costs such that the contestable market equilibrium is feasible and sustainable in each product-destination country market, the household maximizes the CES objective, and the goods and labor markets clear.

Buyer firms in country n compare the average costs $AC_{ni,s}(x^*_{ni,s}(\omega), q_n(\omega))$ of purchasing a quantity $q_n(\omega)$ of product ω across all systems s and origin countries i to minimize overall average costs

$$AC_n(q_n(\omega))^* = \min \left\{ \min \left\{ AC_{ni,A}(x^*_{ni,A}(\omega), q_n(\omega)), AC_{ni,J}(x^*_{ni,J}(\omega), q_n(\omega)) \right\} ;\ i = 1, ..., N \right\}, \quad (15)$$

where $x^*_{ni,s}(\omega)$ is the optimal batch size for each country-system combination determined by the first-order condition (6). As described in the preceding section, since average costs are downward sloping in q and the market is contestable, in equilibrium there is only one buyer firm serving each market. This buyer procures under one system from the seller country with the lowest average costs. The contestable market price is $p_n(\omega) = AC_n(q_n(\omega))^*$.

From the properties of the Cobb-Douglas utility function, households spend a fraction α of their income on manufactured goods. Consumption of each manufactured good is chosen to maximize (12) subject to the budget constraint

$$\int_0^1 p_n(\omega) q_n(\omega)\, d\omega \leq \alpha \left(w_n L_n + \sum_i \sum_s \int \pi^s_{in,s}(\omega) I_{in,s}(\omega) d\omega \right), \quad (16)$$

where $I_{in,s}(\omega)$ is an indicator function that is equal to one if the buyer in country i procures product ω from country n under system s, and zero otherwise. The right-hand side is country n's total income, W_n. As in Eaton and Kortum (2002) and Antras, Fort, and Tintelnot (2017), the wage rate is exogenously pinned down by the productivity of the non-manufacturing sector as $w_n = a_n$. The second term on the right-hand side, which is new relative to Eaton and Kortum (2002), represents the incentive premia collected from shipments to countries i under $s = J$. This term is zero if shipments are under the "American" system. Thus, all else equal, households' income in country n rises with the number of products that are shipped under the "Japanese" system.

Consumption of the non-manufactured good satisfies $Z_n = (1 - \alpha) W_n$. Markets clear for each manufactured good ω, for the non-manufactured good, and labor markets clear in each country. We provide these market clearing conditions in Appendix F.

Due to our departure from competitive markets, our problem does not admit an analytical solution. Below, we solve for the equilibrium numerically using an iteration procedure. First, we trace out the AC curves in each market by solving the cost minimization problem for different values of $q_n(\omega)$. We then guess a price index for manufactured goods p_n and a total income W_n and compute the demand curve for each market from household utility maximization under these choices. Using the intersection of supply and demand, we obtain a candidate equilibrium price vector $p_n(\omega)$, quantity vector $q_n(\omega)$, and a new price index p_n, which determine trade flows, and hence the amount of labor used in each sector and thus total income W_n. We then use the new values of p_n and W_n to obtain new demand curves and iterate to convergence. Appendix G provides further details on our solution algorithm.

7 Quantitative Analysis

7.1 Parametrization and Calibrated Parameters

We now estimate the model quantitatively to analyze the effects of trade policy on trade and aggregate welfare.

Parametrization We set $N = 3$ countries and interpret these countries to be the United States, China, and the Rest of the World (RoW).[44] We assume inspection costs for domestic procurement to be zero, implying that all domestic sourcing takes place under the "American" system. For imports, we assume that the distribution of inspection costs, $G(m)$, is, like that for productivity, Frechet

$$G(m) = e^{-\Theta m^{-\gamma}}, \tag{17}$$

with the parameters to be estimated. We set each time period to one quarter.

Calibrated Parameters We follow Caliendo, Dvorkin, and Parro (2019) in setting the interest rate, r, to be 0.01, and use Antras, Fort, and Tintelnot (2017)'s estimate of the elasticity of substitution, $\sigma = 3.85$. We normalize $\underline{\theta} = 0$ and choose $\bar{\theta} = 1$ for the cost of high quality. This parameter only appears via the ratio $\bar{\theta}/\Upsilon_i(\omega)$ and we will estimate the mean of the distribution of Υ below to match the data. We set the share of consumption accounted for by manufactured goods to $\alpha = 0.5$ to match the share spent on manufactured goods and food in developed countries from Duarte (2020). We follow Eaton and Kortum (2002) in fixing the dispersion of productivity to be the same in each country and equal to $\zeta = 3.6$.

The wages of each country, w_n, are exogenous in the model and pinned down by the productivity of the non-manufacturing sector. We normalize the U.S. to $w_{US} = 1$. Since comparable cross-national wage data are difficult to obtain, we estimate each country's wage to be two-thirds of their GDP divided by the size of their labor force using data reported by the World Bank World Development Indicators (WDI) in 2016. For the Rest of the World, we take a weighted average across countries using each country's exports to the United States in 2016 as weights. This procedure yields $w_{CN} = 0.12$ and $w_{RoW} = 0.47$. We normalize the U.S. labor force $L_{US} = 1$, and set labor for China so that we match the size of the labor force in 2016 from the WDI, yielding $L_{CN} = 5$. We set the labor force for the rest of the world to match the combined workforce of the United States' ten largest trading partners (listed in Table 3), which results in $L_{RoW} = 2.5$. In the model, countries' labor force mainly drives their total income and therefore their imports from the United States. However, it does not significantly affect U.S. imports from these countries, our main object of interest.

We assume that the trade war shocks are symmetric between country pairs, $\rho_{ni} = \rho_{in}$, and that trade wars between the United States and the RoW and China and RoW are unlikely in steady state by setting $\rho_{US,RoW} = 0$ and $\rho_{CN,RoW} = 0$. Nearly all U.S. imports from RoW from 1993 to 2016 were from WTO members, which are subject to a formal dispute settlement system to avoid a trade war, and hence a trade war with these countries would have been seen as very unlikely.

[44]While our model generalizes to an arbitrary number of countries, for our purposes three are sufficient.

To calibrate the probability of a trade war between the U.S. and China, we proceed as follows. First, we take all buyer-seller ($mxhcz$) quintuples in the data that appear to be associated with "Japanese" procurement, We identify these similarly to Section 4 as all quintuples where $mhcz$ is in the first quartile of the SPS distribution in the entire dataset. For these relationships, we compute the separation hazard rate, that is, the probability that a quintuple trades for the last time at age t quarters conditional on having lasted that long, separately for U.S.-China and for U.S.-RoW trade.[45] We then fit the exponential decay function $e^{-\psi_{ni}t}$ to minimize the squared deviation between this function and the empirically observed hazard rate for each country pair. We fit this function from quarter two of the relationship onwards since a large share of quintuples break up after only one transaction for exogenous reasons. We thus obtain $\psi_{US,RoW} = 0.0873$ and $\psi_{US,CN} = 0.1137$. Given our assumption of a zero probability of trade wars between the U.S. and the RoW, we interpret the hazard rate of breakups with the rest of the world as normal churn due to firm exits, product obsolescence, and so on. Given this baseline separation rate, we take the excess probability of break-ups with Chinese suppliers as reflecting the added probability of a trade war, and hence set $\rho_{US,CN} = 0.0264$. Table 10 summarizes the calibrated parameters.

Table 10: Calibrated Parameters

Parameter	Value	Source
Interest rate (r)	0.01	Caliendo, Dvorkin, and Parro (2019)
Elasticity of substitution (σ)	3.85	Antras, Fort, and Tintelnot (2017)
Cost of low quality ($\underline{\theta}$)	0	Normalization
Cost of high quality ($\bar{\theta}$)	1	Normalization
Consumption share of manufactured goods (α)	0.5	Duarte (2020)
Dispersion of productivities (ζ)	3.6	Eaton and Kortum (2002)
Wage (w_i)		
- U.S.	1	Normalization
- China	0.12	World Bank, authors' calculations
- RoW	0.47	World Bank, authors' calculations
Labor Force (L_i)		
- U.S.	1	Normalization
- China	5	World Bank, authors' calculations
- RoW	2.5	World Bank, authors' calculations
Rate of trade wars, U.S.-China ($\rho_{US,CN}$)	0.0264	Census Bureau (LFTTD)

7.2 Targeted Moments and Estimation

Three sets of parameters remain to be estimated: the productivity scales T_n, the country-specific fixed costs f_n, and the inspection cost parameters Θ and γ. We estimate these parameters via a simulated method of moments procedure using moments observed in the LFTTD as well as other datasets. While the parameters are jointly estimated, we proceed to describe the empirical moments targeted and the underlying identification assumptions for each parameter in turn. The column labeled "Data"

[45] We exclude 2016 from this computation due to the censoring of our data.

in Table 11 summarizes the targeted moments and their values in the data. We provide more details on how each moment is constructed in Appendix H.

We normalize $T_{US} = 1$, and estimate the other two parameters using the United States' import penetration from China and from the rest of the world in 2016 (rows 1 and 2 of Table 11). A lower value of T_n increases country n's productivity, which raises that country's share in U.S. consumption and hence imports.

Table 11: Estimated Parameters and Targeted Moments

	Parameter	Value Used in Simulations	Moment Used to Derive Value	Moment in Data	Moment in Model
(1)	Productivity China (T_{CN})	14.498	Share of Chinese imports in consumption	0.074	0.098
(2)	Productivity RoW (T_{RoW})	2.550	Share of RoW imports in consumption	0.270	0.299
(3)	Fixed costs, China (f_{CN})	0.018	$WBS_{A,CN}$	39.07	35.124
(4)	Fixed costs, RoW (f_{RoW})	0.012	$WBS_{A,RoW}$	42.57	46.821
(5)	Scale of inspection costs (Θ)	9.298	WBS difference Q4-Q1, China	0.871	1.466
(6)			WBS difference Q4-Q1, RoW	0.822	0.810
(7)	Dispersion of inspection costs (γ)	0.122	Sd of WBS, China	0.390	0.323
(8)			Sd of WBS, RoW	0.398	0.291

Source: LFTTD and authors' calculations. Column 1 lists parameters needed for the model. Column 2 contains the value of the parameter used our our simulations. Column 3 reports the moment we use to derive this value. Column 4 presents the calculated moment in the data. Column 5 lists the moment calculated from simulations of the model.

The country-specific fixed costs f_n are estimated using the frequency of shipments from China and the RoW to the United States. Proposition 2.2 implies that higher fixed costs lead to shipments that are more spaced out within a given system. Since we do not observe the procurement system with certainty in the data, we focus on buyer quadruples that fall into the fourth quartile of the suppliers per shipments (SPS_{mhcz}) distribution within a given region (China or RoW) by product by mode triple to isolate quadruples that most likely use the "American" system. We then compute the average weeks between shipments (WBS_{mhcz}) across these quadruples (rows 3 and 4), and construct the same moments in the model.

The scale parameter of inspection costs, Θ, governs the frequency of shipments under the "American" system relative to shipments under the "Japanese" system. A lower Θ raises the average inspection cost, making shipments relatively less frequent under "American" procurement by Proposition 2.2. We estimate this moment by running the classification regression (8) with weeks between shipments (WBS_{mhcz}) as dependent variable for China and for the rest of the world,

$$ln(\overline{WBS}_{mhcz}) = \beta_0 + \beta_1 \mathbb{I}_{SPS_{mhcz}=Q4} + \beta_2 \ln(QPW_{mhcz}) + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{hcz} + \epsilon_{mhcz}, \tag{18}$$

where we replace SPS_{mhcz} on the right-hand side with a dummy, $\mathbb{I}_{SPS=Q4}$, indicating whether

SPS_{mhcz} falls into the fourth quartile of its distribution within a given region by product by mode bin. We keep only the first and the fourth quartile of the SPS distribution for the regression to maximize the likelihood of distinguishing between the two systems. A higher estimated coefficient on the dummy indicates that shipments are relatively more dispersed under the "American" system than under "Japanese" sourcing. Rows 5 and 6 of Table 10 present the estimated coefficients.

Finally, the parameter governing the dispersion of inspection costs, γ, is estimated from the dispersion in shipping times across $mhcz$ quadruples. When gamma is low, the inspection cost draws are more dispersed, leading to a higher variance of the shipping frequencies. We construct this moment by running regression (18) with country-mode fixed effects to preserve variation across products. We then keep only $mhcz$ quadruples that fall into the fourth quartile of the SPS distribution, and compute for these observations the standard deviation of the regression residuals across product-country-mode bins. Rows 7 and 8 show the estimated moments.

Our estimation algorithm is standard: we solve for a vector of parameters satisfying

$$\phi^* = \arg\min_{\phi \in \mathbb{F}} \sum_x T(m_x(\phi), \hat{m}_x) \tag{19}$$

where $T(\cdot)$ is the percentage difference between the model, $m_x(\phi)$, and data, $\hat{m}_x$, moments. Appendix I provides more details on the estimation algorithm and outcomes.

7.3 Discussion

Our estimation relies upon a number of assumptions. First, we assume that inspection costs vary across products. This assumption generates dispersion in the relative costs of American- and Japanese-style procurement and thus variation in the procurement system used across goods coming from the same country, since Table 3 suggests that both systems are used by each exporter country. While the dispersion of productivity across products could in principle generate some variability in shipping systems by affecting the relative demand for each product, and hence the relative importance of fixed and variable costs, in practice this mechanism has only a minor effect. Our assumption implies that we need to parametrize the shape of the distribution of inspection costs. This distribution determines the relative costs of "Japanese" versus "American" procurement and hence the welfare costs of changing the probability of a trade war. Importantly, we need to know not only the inspection costs of products that end up being sourced under the "American" system, but also of products that are not sourced via the "American" system.[46] We choose a Frechet distribution to mirror the assumptions on productivity, and fix the shape of the distribution using the dispersion in shipping times of those products that are most likely procured under the "American" system. This distributional with respect to products observed under the "American" system then pins down the inspection costs for all products.[47]

Second, we do not allow the probability of a trade war to vary across products, only across

[46]As an example, nuclear reactors may never be sourced via American-style procurement, but for our counterfactuals we need to know what the costs under the "American" system would have been to compute the welfare costs of shutting down "Japanese" sourcing.

[47]We obtain similar results using a Pareto distribution.

countries. While such variation might be relevant in firms' choice of procurement systems in practice, we cannot disentangle its impact from that of inspection costs without further information on which procurement system firms are using in the data. As a result, we load all of the heterogeneity of procurement system choice onto inspection costs. Similarly, inspection costs vary by product but not by country.

Third, in estimating moments specific to each system, we assume buyer quadruples in the first quartile of the SPS_{mhcz} distribution are "Japanese" and those in the fourth quartile are "American". We construct moments in the model the same way, even though we *do* know which system is chosen there. We then use the estimated model to infer structurally the shares of each procurement system.

Fourth, our model assumes away important sources of heterogeneity present in the data, such as variation in mode of transportation. To the extent possible, we try to account for such heterogeneity in our estimations, e.g., by using product-country-mode fixed effects. We also control for the quadruples' overall quantity traded per week in recognition of general equilibrium effects across estimation runs.

7.4 Model Fit and Results

The column labeled "Value" in Table 11 presents the estimated values of the parameters, and the "Model" column shows the values of the targeted moments. The model provides a good fit along several dimensions. First, it matches the trade shares and generates shipping frequencies consistent with the data. It also generates substantial variation in shipping frequencies across goods, although our model slightly undershoots the empirically observed heterogeneity. This is not surprising: In the data, there are reasons for differences in shipping times across products other than inspection costs, such as seasonality of the good, characteristics of buyers and sellers involved, and the ports most frequently used for a given good. However, overall, our model-generated moments are relatively close to their empirical analogues.

The first column of Table 12 presents some summary statistics of U.S. imports in the estimated equilibrium. The first four rows show the share of manufactured goods consumption that is imported from China and the rest of the world, as well as the share of the imports that are obtained under the "Japanese" system. Our estimates imply that 69 percent of imports from China are under the "Japanese" system, while 76 percent of imports from the rest of the world take place under "Japanese" procurement. The higher share for the rest of the world reflects the higher trade war probability with China, which discourages trade under the "Japanese" system. The structurally estimated "Japanese" shares are in the ballpark of the empirical estimates we obtained using shipments in the first quartile of the SPS_{mhcz} distribution in Table 3, providing an ex-post check of our estimates. Overall, our model suggests that "Japanese" procurement accounts for the majority of imports by value.

Row 5 presents the price index in manufacturing in the United States, $p_{U.S.}$, and row 6 shows the utility $Q_{U.S.}^{\alpha} Z_{U.S.}^{1-\alpha}$. Both of these are normalized to 1 for ease of interpretation. Row 7 shows that the share of labor in the manufacturing sector is 81%, with the rest of labor used to make the non-manufactured good.

Our estimated inspection cost parameters imply that the average product imported by the U.S. is subject to an inspection cost of 0.13 percent of the import value. We estimate fixed costs of 1.93

percent of the import value. These figures are in line with estimates by Kropf and Sauré (2014), who estimate that Swiss exporters face total fixed shipment costs of 0.8 percent to 5.4 percent of the value imported, providing an external validity check. Our method of computing the fixed costs of shipments is similar to theirs, using the shipping frequency and a structural model.

Table 12: Comparison of Equilibria

		Baseline Equilibrium	**Equilibrium Without Japanese Sourcing**
(1)	Value imported from China (%)	9.7%	6.5%
(2)	- of which, "Japanese"	68.9%	.
(3)	Value imported from ROW (%)	29.9%	16.1%
(4)	- of which, "Japanese"	76.1%	.
(5)	Manufacturing price index	1.000	1.065
(6)	Utility	1.000	0.966
(7)	Labor in Manufacturing	0.809	0.787
(8)	Avg. inspection costs	0.13%	2.15%
(9)	Avg. fixed costs	1.93%	0.74%

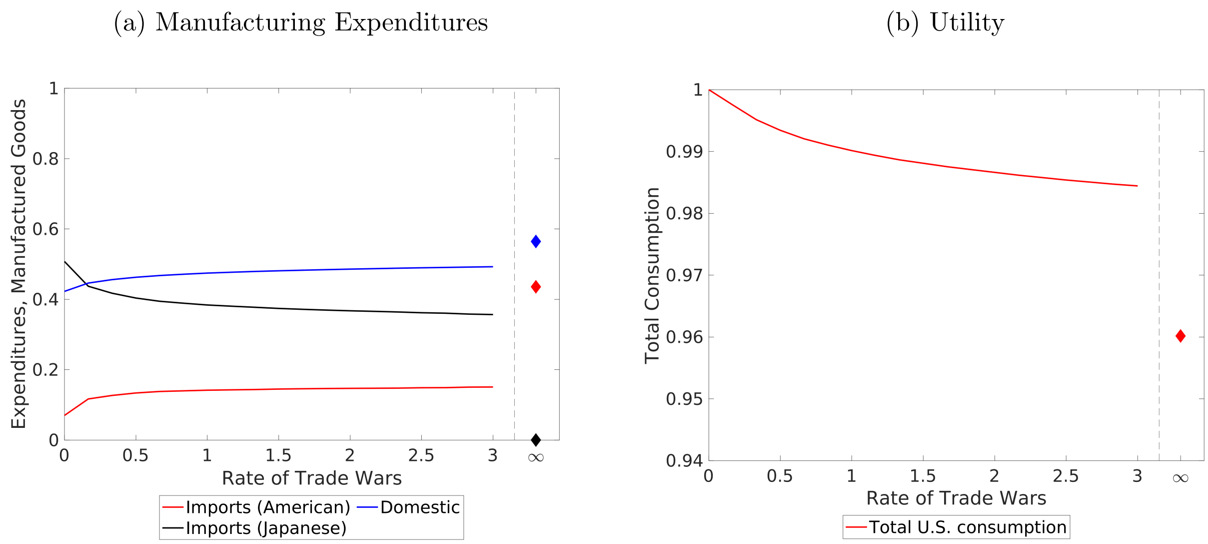
7.5 Counterfactuals

We next analyze the effect of the arrival rate of trade wars on a country's choice of procurement system and source country. The choice of country and system is shaped by three factors: (i) productivity, (ii) the product-specific inspection costs, (iii) and the country-pair-specific probability of trade peace. Products in which the domestic country has a higher productivity than foreign countries are sourced domestically, since domestic sourcing avoids any cost of inspection. Products in which the domestic country is not the most productive and inspection costs are low tend to be sourced under the "American" system. Products in which the domestic country is not the most productive and inspection costs are high are sourced under the "Japanese" system if the trade war arrival rate is sufficiently low. A higher trade war arrival rate reduces the share of goods that are imported by raising the incentive premium for importing, making domestic sourcing more attractive. As the arrival rate of trade wars goes to infinity, no goods are imported under the "Japanese" system since the incentive premium becomes prohibitive.

To illustrate these mechanisms, Figure 5a shows the value of manufactured goods purchased by origin and system at the estimated parameters as $\rho_{US,RoW}$ increases from zero to infinity while keeping the arrival rate of a trade war with China at $\rho_{US,RoW}+0.0264$.[48] As the trade war arrival rate rises, the incentive premium of "Japanese" sourcing increases, reducing imports under the "Japanese" system. Buyers switch to sourcing under the "American" system and to domestic sourcing, leading to an increase in expenditures for these two sources. In general, buyers switch to "American" sourcing for goods where the productivity advantage of importing is sufficiently great to outweigh the increase in

[48] We normalize the total expenditures of manufactured goods at a trade war arrival rate of zero to one.

Figure 5: Effect of Trade War Arrival Rate on Sourcing and Consumption



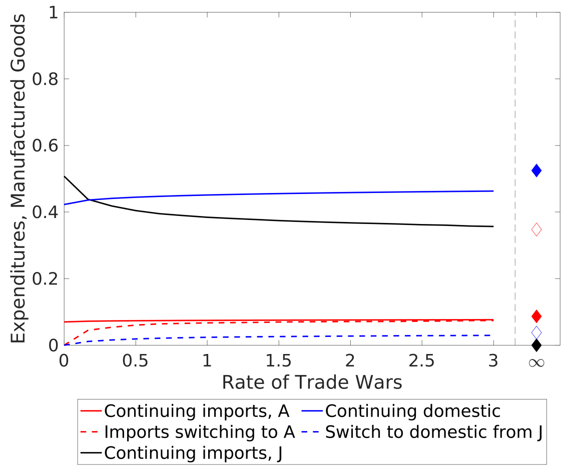
sourcing costs due to having to pay the inspection costs. If the productivity advantage is small or inspection costs are high, buyers switch to domestic sourcing.

The increase in the trade war arrival rate increases sourcing costs, which raises prices and lowers consumption. Figure 5b shows the total utility of U.S. consumers. As the arrival rate of trade wars rises, overall consumption declines and consumer welfare falls.

The second column of Table 12 presents some statistics for our estimated economy when "Japanese" procurement is not possible, i.e., $\rho_{US,RoW} = \infty$. The share of consumption imported from China falls from 9.7 percent to 6.5 percent (row 1), and sourcing from the rest of the world declines by 14 percent percentage points to 16 percent (row 3). The manufacturing price index p_{US} rises due to the higher sourcing costs by 6.5 percent (row 5), causing a decline in consumer welfare by 3.4 percent (row 6). This result highlights that non-traditional trade policies on their own can generate welfare effects by changing the probability of maintaining long-term relationships, even without the presence of fundamental differences in productivity. While some manufacturing production is brought back to the United States, world production declines due to the less efficient production, and U.S. labor in manufacturing falls by 2 percentage points due to the lower world demand (row 7). Finally, since the products which switch from "Japanese" to "American" imports have relatively high inspection costs, the average inspection cost of imported products rises (row 8). However, the average fixed cost as a share of import value falls because import prices have increased (row 9).

Figure 6 further decomposes the change in manufacturing expenditures relative to the baseline into different adjustment margins. First, the solid red line plots the level of expenditures on imports under the "American" system for goods that were imported under the "American" system at $\rho = 0$. We refer to these imports as imports that continue to be American. The dashed red line shows imports that were not under the "American" system at $\rho = 0$ but that are "American" for greater trade war

Figure 6: Decomposition of Adjustment Margins



arrival rates, i.e., products that switch towards the "American" system. Both of these lines increase with the rate of trade wars. The majority of the increase in imports under the "American" system comes from products that switch systems. However, there is also a small increase for continuing goods under the "American" system. Consumers increase their expenditures on these goods because they become relatively cheaper compared to products sourced under the "Japanese" system, and hence there is a substitution effect towards them.

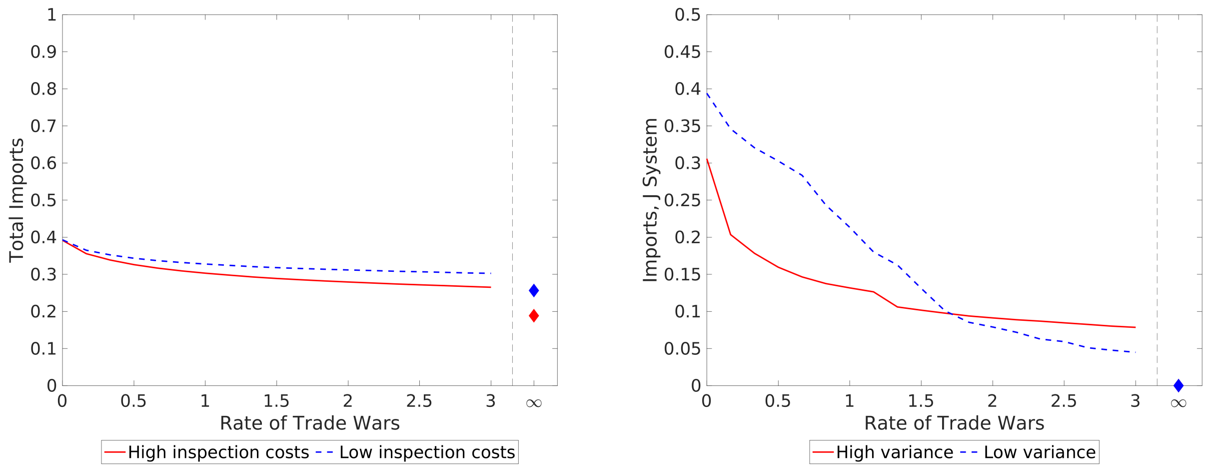
The blue lines decompose the domestic sourcing, for goods that were were already sourced domestically at $\rho = 0$ (solid blue line) and for goods that were sourced via Japanese-style imports at $\rho = 0$ and switch to domestic (dashed blue line). Both lines increase with the arrival rate of trade wars. However, only relatively few goods switch from importing to domestic procurement. Instead, most of the increase in domestic sourcing comes from products that were already obtained domestically at $\rho = 0$, since these products see an increase in demand due to a substitution effect.

The effect of changes in the probability of trade peace on trade depends crucially on the distributions of inspection costs. To illustrate this point, we start from the baseline equilibrium but assume an inspection cost parameter $\Theta = 1$, which increases the average inspection costs relative to the baseline. The red solid line in Figure 7a shows total U.S. imports for different trade war arrival rates under this scenario. For comparison, the blue dashed line presents U.S. imports when average inspection costs are low, $\Theta = 100$. In the high inspection cost scenario, total imports decline relatively quickly with the arrival rate of trade wars. In this case, a worsening in the trade policy environment raises the costs of the "Japanese" system, and since inspections under the "American" system are expensive total imports fall. In contrast, the relationship between imports and trade wars is relatively flat when inspection costs are low. In this case, the "American" system can be used to circumvent the low quality of institutions, which prohibit trade under the "Japanese" system.

The distribution of inspection costs γ across products is similarly important. The red solid line in Figure 7b shows how a change in the trade war arrival rate affects U.S. imports under the "Japanese"

Figure 7: Effect of Distribution of Inspection Costs on Sourcing

(a) Mean Parameter (Θ) (b) Dispersion Parameter (γ)



system when inspection costs are very dispersed across products, $\zeta = 0.2$. In that case, an increase in the arrival rate of trade wars decreases "Japanese" procurement only slightly because only a few additional products become competitive under the "American" system because their inspection costs are so dispersed. The blue dashed line instead shows the analogous effects when the dispersion of inspection costs is small, $\zeta = 2$. In this case, nearly all products have the same inspection costs. As a result, when the trade war arrival rate is low nearly all products are sourced under the "Japanese" system, and when the arrival rate is high all products are sourced under the "American" system, leading to a steep drop in "Japanese" procurement as the trade war arrival rate rises.

8 Conclusion

This paper analyzes the impact of changes in trade policy on procurement patterns along a supply chain using theory, data and quantitative methods. We develop a theoretical model in which importers' ability to solve a quality control problem depends upon exporters' beliefs about the possibility of a trade war breaking out between the firms' countries. When the probability of trade peace is small, buyers choose American-style procurement, characterized by competitive bidding for large, infrequent orders, and costly inspections to ensure the provision of high-quality goods. When the probability of trade peace is high, buyers can induce sellers to provide high quality without inspections by paying them a premium above their costs over a long-term relationship. We show that changes in trade policy that reduce the likelihood of trade wars increase welfare by lowering procurement costs.

We examine the model's key implications using transaction-level U.S. import data. We begin by classifying importer-exporter relationships as American- or Japanese-style and show that these relationships differ along the dimensions – such as shipment size, shipment frequency and shipment size – emphasized in the model. Next we estimate the effect of the U.S. granting of Permanent Normal Trade Relations – which substantially reduced the possibility of a U.S.-China trade war – on the procurement patterns of U.S.-based firms. Using triple difference-in-differences specification, we show that PNTR is associated with a movement toward more Japanese-style procurement among U.S. importers and Chinese exporters along the dimensions highlighted by the model.

Our findings suggest that an important but under-examined aspect of trade agreements in a world with already low tariffs may be their effect on relationship formation. That is, trade agreements promoting institutions that allow firms to develop more stable relationships may give rise to an additional source of welfare gains from trade associated with reducing inventory and monitoring costs.[49] The extent to which such gains are smaller or larger than those that allow firms better access to contract enforcement or dispute resolution is an interesting area for further research.

[49] Indeed, improving the efficiency of trade relationships is a goal of the recent WTO agreement on trade facilitation. See https://www.wto.org/english/thewto_e/minist_e/mc9_e/desci36_e.htm.

References

Abreu, D., D. Pearce, and E. Stacchetti (1990): "Toward a Theory of Discounted Repeated Games with Imperfect Monitoring," *Econometrica: Journal of the Econometric Society*, pp. 1041–1063.

Alessandria, G., J. P. Kaboski, and V. Midrigan (2010): "Inventories, Lumpy Trade, and Large Devaluations," *American Economic Review*, 100(5), 2304–39.

——— (2011): "US Trade and Inventory Dynamics," *American Economic Review*, 101(3), 303–07.

Alessandria, G. A., S. Y. Khan, and A. Khederlarian (2019): "Taking Stock of Trade Policy Uncertainty: Evidence from China's Pre-WTO Accession," Working Paper 25965, National Bureau of Economic Research.

Alviarez, V., M. Fioretti, K. Kikkawa, and M. Morlacco (2021): "Two-Sided Market Power in Firm-to-Firm Trade," .

Amiti, M., M. Dai, R. C. Feenstra, and J. Romalis (2020): "How Did China's WTO Entry Affect US Prices?," *Journal of International Economics*, 126, 103339.

Amiti, M., S. J. Redding, and D. E. Weinstein (2019): "The Impact of the 2018 Tariffs on Prices and Welfare," *Journal of Economic Perspectives*, 33(4), 187–210.

Antoniades, A. (2015): "Heterogeneous Firms, Quality, and Trade," *Journal of International Economics*, 95(2), 263 – 273.

Antràs, P. (2003): "Firms, Contracts, and Trade Structure," *Quarterly Journal of Economics*, 118(4), 1375–1418.

——— (2005): "Incomplete Contracts and the Product Cycle," *American Economic Review*, 95(4), 1054–1073.

Antràs, P., and D. Chor (2018): "Chapter 5 - On the Measurement of Upstreamness and Downstreamness in Global Value Chains," World Trade Evolution: Growth, Productivity and Employment. Routledge.

Antras, P., and D. Chor (2021): "Global Value Chains," .

Antràs, P., T. C. Fort, and F. Tintelnot (2017): "The Margins of Global Sourcing: Theory and Evidence from U.S. Firms," *American Economic Review*, 107(9), 2514–64.

Antras, P., T. C. Fort, and F. Tintelnot (2017): "The Margins of Global Sourcing: Theory and Evidence from US Firms," *American Economic Review*, 107(9), 2514–64.

Antràs, P., and E. Helpman (2008): "Contractual Frictions and Global Sourcing," in *The Organization of Firms in a Global Economy*, ed. by E. Helpman, T. Verdier, and D. Marin, pp. 9–54. Harvard University Press.

Bajari, P., S. Houghton, and S. Tadelis (2014): "Bidding for Incomplete Contracts: An Empirical Analysis of Adaptation Costs," *American Economic Review*, 104(4), 1288–1319.

Baumol, W., J. Panzar, and R. Willig (1982): *Contestable Markets and the Theory of Industry Structure*. Harcourt Brace Jovanovich, New York.

Békés, G., L. Fontagné, B. Muraközy, and V. Vicard (2017): "Shipment frequency of exporters and demand uncertainty," *Review of World Economics*, 153(4), 1610–2886.

Bernard, A. B., J. B. Jensen, and P. K. Schott (2009): "Importers, Exporters and Multinationals: A Portrait of Firms in the U.S. that Trade Goods," in *Producer Dynamics: New Evidence from Micro Data*, ed. by T. Dunne, J. B. Jensen, and M. J. Roberts. Chicago University Press, Chicago.

Boehm, C. E., A. Flaaen, and N. Pandalai-Nayar (2019): "Input Linkages and the Transmission of Shocks: Firm-Level Evidence from the 2011 Tōhoku Earthquake," *The Review of Economics and Statistics*, 101(1), 60–75.

——— (2020): "Multinationals, Offshoring, and the Decline of U.S. Manufacturing," *Journal of International Economics*, 127, 103391.

Bown, C., P. Conconi, A. Erbahar, and L. Trimarchi (2020): "Trade Protection Along Supply Chains," Discussion paper, Working Paper.

Bustos, P. (2011): "Trade Liberalization, Exports and Technology Upgrading: Evidence on the impact of MERCOSUR on Argentinian Firms," *American Economic Review*, 101(1), 304–340.

Caliendo, L., M. Dvorkin, and F. Parro (2019): "Trade and labor market dynamics: General equilibrium analysis of the china trade shock," *Econometrica*, 87(3), 741–835.

Cicala, S. (2015): "When does Regulation Distort Costs? Lessons from Fuel Procurement in U.S. Electricity Generation," *American Economic Review*, 105(1), 411–444.

De Loecker, J., and P. K. Goldberg (2014): "Firm Performance in a Global Market," *Annual Review of Economics*, 6, 201–27.

De Loecker, J., P. K. Goldberg, A. Khandelwal, and N. Pavcnik (2016): "Prices, Markups, and Trade Reform," *Econometrica*, 84(2), 445–510.

Defever, F., C. Fischer, and J. Suedekum (2016): "Relational Contracts and Supplier Turnover in the Global Economy," *Journal of International Economies*, 103, 147–165.

Duarte, M. (2020): "Manufacturing consumption, relative prices, and productivity," *Journal of Macroeconomics*, 65, 103232.

Dyer, J. H., and W. G. Ouchi (1993): "Japanese-Style Partnerships: Giving Companies a Competitive Edge," *Sloan Management Review*, 35(1), 51–63.

Eaton, J., and S. Kortum (2002): "Technology, Geography, and Trade," *Econometrica*, 70(5), 1741–1779.

Fajgelbaum, P. D., P. K. Goldberg, P. J. Kennedy, and A. K. Khandelwal (2019): "The Return to Protectionism*," *The Quarterly Journal of Economics*, 135(1), 1–55.

Feenstra, R. C., and G. H. Hanson (2005): "Ownership and Control in Outsourcing to China: Estimating the Property-Rights Theory of the Firm," *Quarterly Journal of Economics*, 120(2), 729–761.

Feenstra, R. C., J. Romalis, and P. K. Schott (2002): "U.S. Imports, Exports, and Tariff Data, 1989-2001," NBER Working Paper No. 9387.

Fisman, R., and Y. Wang (2010): "Trading Favors within Chinese Business Groups," *American Economic Review: Papers & Proceedings*, 100(2), 429–433.

Flaaen, A., F. Haberkorn, L. Lewis, A. Monken, J. Pierce, R. Rhodes, and M. Yi (2019): "Bill of Lading Data in International Trade Research with an Application to the COVID-19 Pandemic," *Federal Reserve Finance and Economics Discussion Paper*, 2019-086.

Flaaen, A., A. Hortaçsu, and F. Tintelnot (2020): "The Production Relocation and Price Effects of US Trade Policy: The Case of Washing Machines," *American Economic Review*, 110(7), 2103–27.

Flaaen, A., and J. Pierce (2019): "Disentangling the Effects of the 2018-2019 Tariffs on a Globally Connected U.S. Manufacturing Sector," *Federal Reserve Finance and Economics Discussion Paper*, 2019-086.

Goldberg, P. K., and N. Pavcnik (2016): "The Effects of Trade Policy," mineo.

Green, E. J., and R. H. Porter (1984): "Noncooperative Collusion Under Imperfect Price Information," *Econometrica*, pp. 87–100.

Grossman, G. M., and E. Helpman (2004): "Managerial Incentives and International Organization of Production," *Journal of International Economics*, 63(2), 237–262.

Hallak, J. C., and P. K. Schott (2011): "Estimating Cross-Country Differences in Product Quality," *Quarterly Journal of Economics*, 126(1), 417–474.

Han, S.-L., D. P. Wilson, and P. Dant, Shirish (1993): "Buyer-Supplier Relationships Today," *Industrial Marketing Management*, 22(4), 331–338.

Handley, K. (2014): "Exporting under Trade Policy Uncertainty: Theory and Evidence," *Journal of International Economics*, 94(1), 50–66.

Handley, K., and N. Limão (2017): "Policy Uncertainty, Trade, and Welfare: Theory and Evidence for China and the United States," *American Economic Review*, 107(9), 2731–83.

Heide, J. B., and G. John (1990): “Alliances in Industrial Purchasing: The Determinants of Joint Action in Buyer-Supplier Relationships,” *Journal of Marketing Research*, 27(1).

Heise, S. (2019): “Firm-to-firm relationships and the Pass-Through of Shocks: Theory and Evidence,” *FRB of New York Staff Report*, (896).

Helper, S. R., and M. Sako (1995): “Supplier Relations in Japan and in the United States: Are They Converging?,” *Sloan Management Review*, 36(3), 77–84.

Hornok, C., and M. Koren (2015a): “Administrative Barriers to Trade,” *Journal of International Economics*, 96, S110–S122.

——— (2015b): “Per-Shipment Costs and the Lumpiness of International Trade,” *Review of Economics and Statistics*, 97(2), 525–530.

Hummels, D., and A. Skiba (2004): “Shipping the Good Apples Out? An Empirical Confirmation of the Alchian-Allen Conjecture,” *Journal of Political Economy*, 112(6), 1384–1402.

Huneeus, F. (2018): “Production Network Dynamics and the Propagation of Shocks,” *Graduate thesis, Princeton University, Princeton, NJ.*

Jarosch, G. (2016): “Searching for Job Security and the Consequences of Job Loss,” University of Chicago.

Kamal, F., and R. Monarch (2018): “Identifying Foreign Suppliers in US Import Data,” *Review of International Economics*, 26(1), 117–139.

Kamal, F., and H. Tang (2015): “Relational Trade: Evidence from the United States,” mimeograph.

Kandori, M. (2002): “Introduction to Repeated Games with Private Monitoring,” *Journal of Economic Theory*, 102(1), 1–15.

Khandelwal, A. (2010): “The Long and Short (of) Quality Ladders,” *Review of Economic Studies*, 77(4), 1450–1476.

Khandelwal, A., P. Schott, and S.-J. Wei (2013): “Trade Liberalization and Embedded Institutional Reform: Evidence from Chinese Exporter,” *American Economic Review*, 103(6), 2169–95.

Kikkawa, A. K., G. Magerman, and E. Dhyne (2019): “Imperfect Competition in Firm-to-Firm trade,” *SSRN Working paper No. 3389836.*

Kropf, A., and P. Sauré (2014): “Fixed Costs Per Shipment,” *Journal of International Economics*, 92(1), 166–184.

Kugler, M., and E. Verhoogen (2012): “Prices, Plant Size, and Product Quality,” *Review of Economic Studies*, 79(1), 307–339.

Kukharskyy, B. (2016): "Relational contracts and global sourcing," *Journal of International Economics*, 101(C), 123–147.

Kukharskyy, B., and M. Pflüger (2010): "Relational Contracts and the Economic Well-Being of Nations," IZA Discussion Paper No. 5394.

Liker, J. K., and T. Y. Choi (2004): "Building Deep Supplier Relationships," *Harvard Business Review*, 82(12), 104–113.

Lyons, Thomas F., K., A. R. Krachenberg, and J. W. Henke Jr. (1990): "Mixed Motive Marriages: What's Next for Buyer-Supplier Relations?," *Sloan Management Review*, 31(3), 29–36.

Macchiavello, R., and A. Morjaria (2015): "The Value of Relationships: Evidence from a Supply Shock to Kenyan Rose Exports," *American Economic Review*, 105(9), 2911–45.

——— (2020): "Competition and Relational Contracts in the Rwanda Coffee Chain*," *The Quarterly Journal of Economics*, 136(2), 1089–1143.

Mailath, G. J., and L. Samuelson (2006): *Repeated Games and Reputations: Long-Run Relationships*. Oxford university press.

Manova, K., and Z. Yu (2017): "Multi-product firms and product quality," *Journal of International Economics*, 109, 116 – 137.

Martin, J. (2012): "Markups, Quality, and Transport Costs," *European Economic Review*, 56(4), 777–791.

Melitz, M. J. (2003): "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, 71, 1695–1725.

Monarch, R., and T. Schmidt-Eisenlohr (2017): "Learning and the Value of Trade Relationships," *FRB International Finance Discussion Paper*, (1218).

O'Neal, C. R. (1989): "JIT Procurement and Relationship Marketing," *Industrial Marketing Management*, 18(1), 55–63.

Ossa, R. (2014): "Trade Wars and Trade Talks with Data," *American Economic Review*, 104(12), 4104–4146.

Pierce, J., and P. K. Schott (2016): "The Surprisingly Swift Decline of U.S. Manufacturing Employment," *American Economic Review*, 106(7), 1632–1662.

Pierce, J. R., and P. K. Schott (2012): "Concording U.S. Harmonized System Categories Over Time," *Journal of Official Statistics*, 28(1), 53–68.

Rauch, J. E. (1999): "Networks versus Markets in International Trade," *Journal of International Economics*, 48(1), 7–35.

Spencer, B. J. (2005): "International Outsourcing and Incomplete Contracts," *Canadian Journal of Economics*, 38(4), 1107–1135.

Tadelis, S., and F. Zettelmeyer (2015): "Information Disclosure as a Matching Mechanism: Theory and Evidence from a Field Experiment," *American Economic Review*, 105(2), 886–905.

Taylor, C. R., and S. N. Wiggins (1997): "Competition or Compensation: Supplier Incentives under the American and Japanese Subcontracting Systems," *American Economic Review*, 87(4), 598–618.

Tirole, J. (1988): *The Theory of Industrial Organization*. MIT Press, Cambridge, MA.

Verhoogen, E. A. (2008): "Trade, Quality Upgrading, and Wage Inequality in the Mexican Manufacturing Sector," *The Quarterly Journal of Economics*, 123(2), 489–530.

WorldBank (2019): "Trading for Development in the Age of Global Value Chains, World Development Report 2020," .

Appendix

A Analytical Results

A.1 Proofs

A.1.1 Second Order Conditions Hold

American System The second derivative of the average cost yields

$$AC_A''(x,q) = \frac{r}{q}\frac{\left(\frac{r}{q}\right)e^{-rx/q}\frac{\bar{\theta}}{\Upsilon}\left[-2\left(1-e^{-rx/q}\right)+\left(\frac{r}{q}\right)\left[1+e^{-rx/q}\right]\left[x+\frac{f+m}{\bar{\theta}/\Upsilon}\right]\right]}{\left[1-e^{-rx/q}\right]^3}.$$

Thus the first order condition is strictly upward sloping, $AC_A''(x,q)>0$, if and only if

$$\left[1+e^{-rx/q}\right]\left[r\frac{x}{q}+\left(\frac{r}{q}\right)\left(\frac{f+m}{\bar{\theta}/\Upsilon}\right)\right]-2\left[1-e^{-rx/q}\right]>0. \tag{A.1}$$

Consider the case when $f+m=0$. If the condition holds for this case, it must also hold for $f+m>0$,because (A.1) is increasing in $f+m$. Define $y\equiv rx/q$. Note that for $y=0$ and $f+m=0$ the left-hand side of equation (A.1) is equal to zero. Taking the derivative of the left-hand side of equation (A.1) with respect to y we obtain

$$1-e^{-y}(1-y),$$

Thus, the left-hand side of (A.1) is strictly increasing in y for $0<y<1$. Therefore, if $0<y<1$, then $AC_A''(x,q)>0$.

Japanese System

$$\begin{aligned} AC_J''(x) = \Bigg[& \frac{\left(\frac{r}{q}\right)^2 e^{-rx/q}\left[f+\underline{\theta}\frac{1}{\Upsilon}x+e^{(r+\rho)x/q}(\bar{\theta}-\underline{\theta})\frac{1}{\Upsilon}x\right]\left[1+e^{-rx/q}\right]}{\left[1-e^{-rx/q}\right]^3} \\ & -\frac{2\left(\frac{r}{q}\right)e^{-rx/q}\left[\underline{\theta}\frac{1}{\Upsilon}+e^{(r+\rho)x/q}(\bar{\theta}-\underline{\theta})\frac{1}{\Upsilon}\left(1+\left(\frac{r+\rho}{q}\right)x\right)\right]\left[1-e^{-rx/q}\right]}{\left[1-e^{-rx/q}\right]^3} \\ & +\frac{\left(\frac{r+\rho}{q}\right)e^{(r+\rho)x/q}(\bar{\theta}-\underline{\theta})\frac{1}{\Upsilon}\left[2+\left(\frac{r+\rho}{q}\right)x\right]\left[1-e^{-rx/q}\right]^2}{\left[1-e^{-rx/q}\right]^3}\Bigg]\frac{r}{q}. \end{aligned}$$

Then $AC_J''(x)>0$, if and only if the numerator is greater than zero. Note that the numerator increases in f. Therefore if the numerator is positive for $f=0$, it must also be positive for $f>0$. Assume $f=0$, and factor the numerator of $AC_J''(x)$ to obtain

$$\begin{aligned} & \left(\frac{r}{q}\right)e^{-rx/q}\left[\underline{\theta}\frac{1}{\Upsilon}+e^{(r+\rho)x/q}(\bar{\theta}-\underline{\theta})\frac{1}{\Upsilon}\right]\left[\left(\frac{r}{q}\right)x\left(1+e^{-rx/q}\right)-2\left(1-e^{-rx/q}\right)\right] \\ + & \left(\frac{r+\rho}{q}\right)e^{(r+\rho)x/q}(\bar{\theta}-\underline{\theta})\frac{1}{\Upsilon}\left[1-e^{-rx/q}\right]\left\{\left[1-e^{-rx/q}\right]\left[2+\left(\frac{r+\rho}{q}\right)x\right]-2\left(\frac{r}{q}\right)xe^{-rx/q}\right\} \end{aligned}$$

Define $y\equiv rx/q$. For the first term note that $(1+e^{-y})\,y-2\,(1-e^{-y})>0$ for $0<y<1$. For the second term to be positive, we require that $\left([1-e^{-y}]\left[2+y+\left(\frac{\rho}{q}\right)x\right]-2ye^{-y}\right)>0$. If $\rho=0$, then $(\cdot)>0$ for $0<y<1$.

Because $(\cdot)$increases in ρ, it must be true that $(\cdot) > 0$ for $\rho > 0$ and $0 < y < 1$. Therefore, if $\rho > 0$ and $0 < y < 1$, then $AC''_J(x) > 0$.

A.1.2 Effect of Quality and Arrival Rate of Trade Wars on Average Costs

$$\left.\frac{\partial p_J}{\partial \underline{\theta}}\right|_{\underline{\theta}<\bar{\theta}} = \left[\left(e^{(r+\rho)x_J/q} - 1\right) x_J r\right] \Big/ \left[\Upsilon(e^{-rx/q} - 1)q\right] < 0$$

$$\left.\frac{\partial p_J}{\partial \bar{\theta}}\right|_{\underline{\theta}<\bar{\theta}} = x_J r e^{(r+\rho)x_J/q} \Big/ \left[\Upsilon(e^{-rx_J/q} - 1)q\right] > 0$$

$$\frac{\partial p_J}{\partial \rho} = \left(e^{(r+\rho)x_J/q} x_J^2(\bar{\theta} - \underline{\theta})r)\right) \Big/ q^2\Upsilon\left(1 - e^{-\frac{rx}{q}}\right) > 0$$

Finally, comparing procurement costs in both systems note that:

$$\frac{r}{q}\frac{f + \bar{\theta}\frac{1}{\Upsilon}x_J^* + (\bar{\theta} - \underline{\theta})\frac{1}{\Upsilon}x_J^*\left[e^{rx_J^*/q} - 1\right]}{1 - e^{-rx_J^*/q}} > \frac{r}{q}\frac{f + \bar{\theta}\frac{1}{\Upsilon}x_J^*}{1 - e^{-rx_J^*/q}} > \frac{r}{q}\frac{f + \bar{\theta}\frac{1}{\Upsilon}x_A^*}{1 - e^{-rx_A^*/q}}$$

The first inequality holds since $e^{rx_J^*/q} > 1$, the second inequality holds because the batch size that minimizes average costs in the “Japanese” system is strictly less than the batch size that minimizes average costs in the “American” system when $m = 0$,$x_J^* < x_A^*(m = 0)$. Hence, the average procurement cost under the “Japanese” system is strictly greater than under the “American” system for any $\rho \geq 0$ when $m = 0$.

A.1.3 Proof of Proposition 2.1

For $\bar{\theta} - \underline{\theta} > 0$ and $\rho > 0$, when $m_A = 0$ average costs under the “Japanese” system must be higher than under the “American” system by the discussion above Proposition 2.1 and in Appendix A.1.2. Since average costs under the “American” system grow without bound as $m_A \to \infty$, there must be an m^* such that average costs under the systems are equalized.

A.1.4 Proof of Proposition 2.2

Japanese System: We apply the implicit function theorem to the FOC (6). We obtain

$$\frac{\partial FOC_J}{\partial \rho} = \frac{2xe^{\frac{x\rho}{q}}\left(\bar{\theta} - \underline{\theta}\right)}{q^2\Upsilon(e^{-\frac{rx}{q}} - 1)}\left[\frac{x\rho}{2}\left(e^{\frac{rx}{q}} - 1\right) + q\left(\left(\frac{rx}{2q} + 1\right)e^{\frac{rx}{q}} - \frac{rx}{q} - 1\right)\right]$$

Define $y = rx/q$. Note that $\lim_{y \downarrow 0} \left(\frac{y}{2} + 1\right)e^y - y - 1 = 0$ and $\frac{d}{dy}\left(\frac{y}{2} + 1\right)e^y - y - 1 = -1 + \frac{1}{2}(x + 3)e^x > 0$. Therefore $\frac{\partial FOC_J}{\partial \rho} > 0$. Then by the implicit function theorem

$$\frac{\partial x}{\partial \rho} = -\frac{\frac{\partial FOC_J}{\partial \rho}}{SOC_J} < 0.$$

Remember that $v_J(x_J, \rho) = f + \bar{\theta}\frac{1}{\Upsilon}x_J^* + (\bar{\theta} - \underline{\theta})\frac{1}{\Upsilon}x_J^*\left[e^{rx_J^*/q} - 1\right]$. Average costs in the “Japanese” system are then $\frac{r}{q}\frac{v_J(x_J,\rho)}{1-exp(-\frac{rx_J}{q})}$. Taking the first order condition of these average costs and setting zero we can write.

$$\frac{\partial v(x_J, \rho)}{\partial x_J} = \frac{r}{q}\frac{v(x_J, \rho)exp(-\frac{rx_J}{q})}{1 - exp(-\frac{rx_J}{q})}$$

Now take the derivative of the unit value, $\frac{v_J(x_J,\rho)}{x_J}$,with respect to ρ to obtain

$$\left(\frac{\partial v(x_J,\rho)}{\partial x_J}\frac{\partial x_J}{\partial \rho}x+\frac{\partial v(x_J,\rho)}{\rho}x_J-v(x_J,\rho)\frac{\partial x_J}{\partial \rho}\right)\frac{1}{x_J^2}$$

Substituting for $\frac{\partial}{\partial x}v(x_J,\rho)$ from the equilibrium condition (22) into (23) we can rewrite (23) to obtain

$$\left[\left(\frac{rx_J}{q}\frac{exp(-\frac{rx_J}{q})}{1-exp(-\frac{rx_J}{q})}-1\right)\frac{\partial x_J}{\partial \rho}v(x_J,\rho)+\frac{\partial v(x_J,\rho)}{\rho}x_J\right]\frac{1}{x_J^2}$$

Note that $\frac{\partial v(x_J,\rho)}{\rho}x_J=\frac{x_J^3(\bar{\theta}-\underline{\theta})}{exp(-\frac{(r+\rho)x_J}{q})q\Upsilon}>0$. Also note that $\frac{rx_J}{q}\frac{exp(-\frac{rx_J}{q})}{1-exp(-\frac{rx_J}{q})}-1<0$ for $0<\frac{rx}{q}<1$. Then because $\frac{\partial x_J}{\partial \rho}<0$ we have shown that $\frac{\partial}{\partial \rho}\frac{v_J(x_J,\rho)}{x_J}>0$

American System: We apply the implicit function theorem to show:

$$\frac{\partial x_A^*}{\partial m}=-\frac{\frac{\partial FOC_A}{\partial m}}{SOC_A}=\frac{r^2e^{-\frac{rx_A}{q}}}{q^2\left(1-e^{-\frac{rx_A}{q}}\right)^2}>0$$

Note that unit values in the "American" system are simply $\frac{v_A(x_A)}{x_A}=\frac{f}{x_A}+\frac{\bar{\theta}}{\Upsilon}$.Therefore, $\frac{\partial x_A^*}{\partial m}>0\Rightarrow\frac{\partial\frac{v_A(x_A)}{x_A}}{\partial m}<0$.

A.1.5 Proof of Proposition 2.3

Part 1: Comparing shipping sizes: $x_J^*<x_A^*$ First note that if $m=0$ and $\bar{\theta}-\underline{\theta}=0$, then average costs in the two procurement systems are identical. If $\frac{\partial x_A^*}{\partial m}>0$ and $\frac{\partial x_J^*}{\partial \underline{\theta}}>0$, then $x_J^*<x_A^*$all else equal. We apply the implicit function theorem. Let $FOC_A$and FOC_J denote the first order conditions to minimize average procurement costs, and, let $SOC_A>0$ and $SOC_J>0$ be the associated second order conditions that are greater than zero by (A.1.1).

American System

$$\frac{\partial x_A^*}{\partial m}=-\frac{\frac{\partial FOC_A}{\partial m}}{SOC_A}=\frac{r^2e^{-\frac{rx_A}{q}}}{q^2\left(1-e^{-\frac{rx_A}{q}}\right)^2}>0$$

Japanese System

$$\begin{aligned}\frac{\partial x_J^*}{\partial \underline{\theta}}=-\frac{\frac{\partial FOC_J}{\partial \underline{\theta}}}{SOC_J}&=\left(\frac{r}{q}\right)\frac{1}{\Upsilon}\frac{\left[1-e^{(r+\rho)x_J^*/q}\left[1+\left(\frac{r+\rho}{q}\right)x_J^*\right]\right]\left[1-e^{-rx_J^*/q}\right]}{\left(1-e^{-rx_J^*/q}\right)^2}\\&-\left(\frac{r}{q}\right)^2\frac{1}{\Upsilon}\frac{x_J^*e^{-rx_J^*/q}\left[1-e^{(r+\rho)x_J^*/q}\right]}{\left(1-e^{-rx_J^*/q}\right)^2}.\end{aligned}$$

For $(r+\rho)x_J^*/q>0$, this expression is negative if and only if

$$\frac{\left[1-e^{(r+\rho)x_J^*/q}\left[1+\left(\frac{r+\rho}{q}\right)x_J^*\right]\right]}{\left[1-e^{(r+\rho)x_J^*/q}\right]}>\frac{\left(\frac{r}{q}\right)x_J^*e^{-rx_J^*/q}}{\left[1-e^{-rx_J^*/q}\right]}. \tag{A.2}$$

Note that the left-hand side is greater than 1. Hence, we need to show that the right-hand side is less than 1. Define $y\equiv rx_J^*/q$, where $0<y<1$. We find for the right-hand side $\lim_{y\to 0}\frac{ye^{-y}}{1-e^{-y}}=\lim_{y\to 0}1-y=1$. Next, note that $\dfrac{d}{dy}\dfrac{ye^{-y}}{1-e^{-y}} = \dfrac{e^{-y}\left[(1-y)-e^{-y}\right]}{\left[1-e^{-y}\right]^2} < 0$. It follows that the right-hand side of (A.2) is never greater than 1. Therefore, $\partial FOC/\partial \underline{\theta} < 0$ and$\partial x_J^*/\partial \underline{\theta} > 0$.

Part 2: Comparing unit values: $v_A(x_A)/x_A < v_J(x_J)/x_J$

$$v_s(x_s)/x_s = \begin{cases} \frac{f}{x_A^*} + \frac{\bar{\theta}}{\Upsilon} & \text{if } s = A \\ \frac{f}{x_J^*} + \frac{\bar{\theta}}{\Upsilon} + \left(e^{\frac{(r+\rho)x}{q}} - 1\right)(\bar{\theta} - \underline{\theta})\dfrac{1}{\Upsilon} & \text{if } s = J \end{cases}$$

Comparing the expressions, $x_A^* > x_J^*$(see Part 1) and $\left(e^{\frac{(r+\rho)x}{q}} - 1\right)(\bar{\theta} - \underline{\theta})\dfrac{1}{\Upsilon} \Rightarrow v_A(x_A)/x_A < v_J(x_J)/x_J$.

A.1.6 Proof of Proposition 2.4

Part 1: Order size and shipping frequency increase in q.

American System We apply the implicit function theorem to the first order condition in the "American" system. From the first order condition and setting to zero we obtain $v'(x) = \frac{r(v(x)+m)e^{-rx/q}}{q\left(1-e^{-rx/q}\right)}$. Substituting this optimality condition into $\frac{\partial FOC_A}{\partial q}$we obtain

$$\frac{\partial x_A}{\partial q} = -\frac{\frac{\partial FOC_A}{q}}{SOC_A} = \frac{\left[1 - \frac{\frac{rx}{q}\mathrm{e}^{-\frac{rx}{q}}}{1-\mathrm{e}^{-\frac{rx}{q}}} - \frac{rx}{q}\right]}{SOC_A} \frac{r^2\left(v\left(x\right)+m\right)\mathrm{e}^{-\frac{rx}{q}}}{q^3\left(1-\mathrm{e}^{-\frac{rx}{q}}\right)^2}$$

Then, $0 < \frac{rx}{q} < 1 \Rightarrow [\cdot] < 0 \Rightarrow \frac{\partial x_A}{\partial q} > 0$ over the relevant parameter range where costs are positive.

For the shipment frequency, $d(x_A^*/q)/dq < 0$, define $\psi_A = x_A^*/q$. Then, simplifying the first-order condition under the "American" system we have

$$FOC(\psi_A) = \bar{\theta}\frac{1}{\Upsilon}\left[1 - e^{-r\psi_A}\right] - \left(\frac{r}{q}\right)e^{-r\psi_A}\left[f + m + \bar{\theta}\frac{1}{\Upsilon}q\psi_A\right] = 0.$$

Applying the implicit function theorem to this expression yields

$$\frac{\partial \psi_A}{\partial q} = -\frac{\frac{\partial FOC(\psi_A)}{\partial q}}{\frac{\partial FOC(\psi_A)}{\partial \psi_J}} = -\frac{[f+m]}{rq\left[f + m + \bar{\theta}\frac{1}{\Upsilon}q\psi_A\right]} < 0,$$

and hence the time between shipments decreases, i.e., shipping frequency increases.

Japanese System We follow the same strategy as in the proof for the American system. From the first order condition, FOC_J, we obtain $\frac{\partial v_J(x_J,q)}{\partial x_J} = \frac{rv_J(x_J,q)e^{-\frac{rx}{q}}}{q\left(1-e^{-\frac{rx}{q}}\right)}$ which we substitute into $\frac{\partial FOC_J}{\partial q}$to obtain:

$$\begin{aligned} \frac{\partial FOC_J}{q} &= \left[1 - \frac{rxe^{-\frac{rx}{q}}}{q\left(1-e^{-\frac{rx}{q}}\right)} - \frac{rx}{q}\right]\left(\frac{r^2 v(x,q)e^{-\frac{rx}{q}}}{q^3\left(1-e^{-\frac{rx}{q}}\right)^2}\right) \\ &- \frac{2(r+\rho)(\bar{\theta}-\underline{\theta})xre^{\frac{x\rho}{q}}}{q^4\Upsilon(e^{-\frac{rx}{q}}-1)^2}\left(\frac{x\rho}{2}\left(e^{\frac{rx}{q}}-1\right) + \left[\left(\frac{rx}{2q}+1\right)e^{\frac{rx}{q}} - \frac{rx}{q} - 1\right]q\right) \end{aligned}$$

Note that $0 < \frac{rx}{q} < 1 \Rightarrow \left[1 - \frac{rxe^{-\frac{rx}{q}}}{q\left(1-e^{-\frac{rx}{q}}\right)} - \frac{rx}{q}\right] < 0$ & $\left[\left(\frac{rx}{2q}+1\right)e^{\frac{rx}{q}} - \frac{rx}{q} - 1\right] > 0 \Rightarrow -\frac{\frac{\partial FOC_J}{q}}{SOC_J} > 0 \Rightarrow \frac{\partial x_J^*}{\partial q} > 0$, because all other terms are positive by inspection.

To see that $d(x_J^*/q)/dq < 0$, define $\psi_J = x_J^*/q$. The first-order condition under the "Japanese" system can then be simplified to

$$
\begin{aligned}
FOC(\psi_J) = &\left[\underline{\theta}\tfrac{1}{\Upsilon} + \left(\bar{\theta} - \underline{\theta}\right)\tfrac{1}{\Upsilon}e^{(r+\rho)\psi_J}\left[1 + (r+\rho)\psi_J\right]\right]\left(1 - e^{-r\psi_J}\right) \qquad \text{(A.3)}\\
&- \left(\frac{r}{q}\right)e^{-r\psi_J}\left[f + \underline{\theta}\tfrac{1}{\Upsilon}\psi_J q + (\bar{\theta} - \underline{\theta})\tfrac{1}{\Upsilon}e^{(r+\rho)\psi_J}\psi_J q\right] = 0.
\end{aligned}
$$

Applying the implicit function theorem to this expression yields

$$\frac{\partial \psi_J}{\partial q} = -\frac{\frac{\partial FOC(\psi_J)}{\partial q}}{\frac{\partial FOC(\psi_J)}{\partial \psi_J}}.$$

For the numerator, we have

$$\frac{\partial FOC(\psi_J)}{\partial q} = \frac{r}{q^2}e^{-r\psi_J}f > 0.$$

For the denominator we find

$$
\begin{aligned}
\frac{\partial FOC(\psi_J)}{\partial \psi_J} = &(r+\rho)(\bar{\theta} - \underline{\theta})\tfrac{1}{\Upsilon}e^{(r+\rho)\psi_J}\left[2 + (r+\rho)\psi_J\right]\left[1 - e^{-r\psi_J}\right]\\
&+ \frac{r^2}{q}e^{-r\psi_J}\left[f + \underline{\theta}\tfrac{1}{\Upsilon}\psi_J + (\bar{\theta} - \underline{\theta})\tfrac{1}{\Upsilon}e^{(r+\rho)\psi_J}\psi_J\right] > 0.
\end{aligned}
$$

Therefore, $\partial FOC(\psi_J)/\partial q > 0$, and thus $d(x_J^*/q)/dq < 0$.

A.1.7 Proof of Lemma 2.5: Average cost curves are downward sloping, convex and reach a limit

Part 1: Average cost curves are downward sloping

American System The average cost function under the "American" system is

$$AC(q) = \frac{\theta\frac{x}{q} + \frac{f}{q} + \frac{m}{q}}{1 - exp(-\frac{rx}{q})}.$$

Taking the first derivative of the expression with respect to q, and fully writing out also the terms that involve x, we get

$$AC'(q) = \frac{-\frac{f+m}{q^2} + \theta\frac{x'(q)}{q} - \theta\frac{x}{q^2}}{1 - exp(-\frac{rx}{q})} - \frac{\frac{r}{q}exp(-\frac{rx}{q})\left[\theta\frac{x}{q} + \frac{f}{q} + \frac{m}{q}\right]x'(q)}{\left[1 - exp(-\frac{rx}{q})\right]^2} + \frac{\left(\frac{rx}{q^2}\right)exp(-\frac{rx}{q})\left[\theta\frac{x}{q} + \frac{f}{q} + \frac{m}{q}\right]}{\left[1 - exp(-\frac{rx}{q})\right]^2}.$$

Re-arranging this expression, we obtain

$$AC'(q) = \frac{-\frac{f+m}{q^2}}{1-exp(-\frac{rx}{q})} + \frac{1}{q}x'(q)\left\{\frac{\theta}{1-exp(-\frac{rx}{q})} - \frac{\frac{r}{q}exp(-\frac{rx}{q})\left[\theta x + f + m\right]}{\left[1-exp(-\frac{rx}{q})\right]^2}\right\}$$
$$-\frac{x}{q^2}\left\{\frac{\theta}{1-exp(-\frac{rx}{q})} - \frac{\frac{r}{q}exp(-\frac{rx}{q})\left[\theta x + f + m\right]}{\left[1-exp(-\frac{rx}{q})\right]^2}\right\}.$$

Note that the two terms in brackets are the first-order condition of the cost function with respect to x, which is equal to zero (this is the "Envelope condition")! This is key: because in the average cost function x and q almost always appear as x/q, we can re-arrange terms to not only cancel the expression containing $x'(q)$, but also the term involving x/q^2. Thus, we get

$$AC'(q) = \frac{-\frac{f+m}{q^2}}{1-exp(-\frac{rx}{q})}.$$

Now, it is much easier to work with this! Also, note that this clearly shows that average cost curves are decreasing.

Japanese System The proof proceeds in the same way as before. Average costs under the "Japanese" system are

$$AC(q) = \frac{\theta\frac{x}{q}exp(\frac{(r+\rho)x}{q}) + \frac{f}{q}}{1-exp(-\frac{rx}{q})}.$$

The first derivative with respect to q is (ignoring the derivative with respect to x here, which we know must be zero)

$$AC'(q) = \frac{-\frac{f}{q^2} - \theta\frac{x}{q^2}exp(\frac{(r+\rho)x}{q}) - \theta(r+\rho)\frac{x^2}{q^3}exp(\frac{(r+\rho)x}{q})}{1-exp(-\frac{rx}{q})} + \frac{\left(\frac{rx}{q^2}\right)exp(-\frac{rx}{q})\left[\theta\frac{x}{q}exp(\frac{(r+\rho)x}{q}) + \frac{f}{q}\right]}{\left[1-exp(-\frac{rx}{q})\right]^2}.$$

Re-arranging yields

$$AC'(q) = \frac{-\frac{f}{q^2}}{1-exp(-\frac{rx}{q})} - \frac{x}{q^2}\left\{\frac{\theta exp(\frac{(r+\rho)x}{q})\left[1+(r+\rho)\frac{x}{q}\right]}{1-exp(-\frac{rx}{q})} - \frac{\frac{r}{q}exp(-\frac{rx}{q})\left[\theta x exp(\frac{(r+\rho)x}{q}) + f\right]}{\left[1-exp(-\frac{rx}{q})\right]^2}\right\}.$$

Similar to before, the term in curly brackets is the first-order condition with respect to x and is equal to zero. Therefore, we have

$$AC'(q) = \frac{-\frac{f}{q^2}}{1-exp(-\frac{rx}{q})}.$$

This function must be convex because the function under the American system was convex for all m, and thus also for $m = 0$.

Part 2: Average cost curves are convex

American System From A.1.6 we obtain the slope of the average cost curve in q :

$$AC'(q) = \frac{-\frac{f+m}{q^2}}{1 - exp(-\frac{rx}{q})}.$$

Taking the second derivative of average costs then yields

$$AC''(q) = \frac{2\frac{f+m}{q^3}}{1 - exp(-\frac{rx}{q})} - \frac{\left(\frac{rx}{q^2}\right) exp(-\frac{rx}{q}) \left(\frac{f+m}{q^2}\right)}{\left[1 - exp(-\frac{rx}{q})\right]^2} + \frac{\left(\frac{rx'(q)}{q}\right) exp(-\frac{rx}{q}) \left(\frac{f+m}{q^2}\right)}{\left[1 - exp(-\frac{rx}{q})\right]^2}.$$

The last term is positive since $x'(q) > 0$. Therefore, to prove that the average cost function is convex, we only need to show that the first two terms together are positive. These terms can be re-written as

$$\frac{2\left[1 - exp(-\frac{rx}{q})\right]\left(\frac{f+m}{q^3}\right) - \left(\frac{rx}{q}\right) exp(-\frac{rx}{q}) \left(\frac{f+m}{q^3}\right)}{\left[1 - exp(-\frac{rx}{q})\right]^2},$$

which is positive if

$$2\left[1 - exp(-\frac{rx}{q})\right] > \left(\frac{rx}{q}\right) exp(-\frac{rx}{q}).$$

This expression holds if

$$2\left[exp(\frac{rx}{q}) - 1\right] > \left(\frac{rx}{q}\right),$$

which is true. Therefore, average costs are convex, for any m and f.

Japanese System From A.1.6 we obtain the slope of the average cost curve in the "Japanese" system.

$$AC'(q) = \frac{-\frac{f}{q^2}}{1 - exp(-\frac{rx}{q})}.$$

By the same arguments as in the "American" system $AC''(q) > 0$.

Asymptote for both systems We first show $(x(q)/q) \to 0$ as $q \to \infty$.

From the Monotone Convergence Theorem, since $(x(q)/q)$ is strictly decreasing and bounded from below by zero, it must converge to a limit. Call this limit $\psi^* \geq 0$. To show that $\psi^* = 0$, assume for contradiction that $\psi^* = K > 0$. Then, it must be the case that there exists no combination of $\psi = x(q)/q < K$ and q that solves the first-order condition of the cost minimization problem. Thus, if we can find a q solving the first-order condition for a $\psi < K$, then K cannot have been the limit since ψ is strictly decreasing.

For the "American" system, pick any $0 \leq \psi_A < K$. The first-order condition of the cost minimization problem under the American system is

$$\bar{\theta}\frac{w_z}{\Upsilon}\left[1 - e^{-r\psi_A}\right] = \left(\frac{r}{q}\right) e^{-r\psi_A}\left[f + mw_b + \bar{\theta}\frac{w_z}{\Upsilon} q\psi_A\right].$$

Re-arranging this expression, we can solve for q as a function of ψ_A and find that

$$q = \frac{\left[f + mw_b\right] re^{-r\psi_A}}{\bar{\theta}\frac{w_z}{\Upsilon}\left[1 - e^{-r\psi_A}\left[1 + r\psi_A\right]\right]}. \tag{A.4}$$

This expression gives the q that solves the first-order condition for a given pick of $\psi_A = x_A/q$. If we can show that for any pick $\psi_A \geq 0$ there exists a $q \geq 0$ solving the equation, then it cannot be the case that $K > 0$ is the limit. For this result to hold, we need to show that the denominator is non-negative. To see that it is non-negative, note that

$$\begin{aligned} 1 - e^{-r\psi_A}\left[1 + r\psi_A\right] &\geq 0 \\ \Leftrightarrow \quad e^{r\psi_A} &\geq 1 + r\psi_A, \end{aligned}$$

which holds. Thus, for any $\psi_A \geq 0$ there exists a $q \geq 0$ solving the equation. In particular, such a q exists for any $\psi_A < K$. Therefore, $(x(q)/q)$ must converge to zero. Indeed, from the equation we can see that for $\psi_A = 0$, q must be infinite.

We can construct a similar proof for the "Japanese" system. The first-order condition under the "Japanese" system is

$$\frac{e^{(r+\rho)\psi_J}\bar{\theta}\frac{w}{\Upsilon}\left[1 + (r+\rho)\psi_J\right]}{1 - e^{-r\psi_J}} = \frac{\left(\frac{r}{q}\right) e^{-r\psi_J}\left[f + e^{(r+\rho)\psi_J}\bar{\theta}\frac{w}{\Upsilon}q\psi_J\right]}{\left[1 - e^{-r\psi_J}\right]^2}.$$

We can re-arrange this expression to solve for q and find that

$$q = \frac{fre^{-r\psi_J}}{\bar{\theta}\frac{w_z}{\Upsilon}e^{(r+\rho)\psi_J}\left[(r+\rho)\psi_J\left[1 - e^{-r\psi_J}\right] + 1 - e^{-r\psi_J}\left[1 + r\psi_J\right]\right]}. \tag{A.5}$$

By the same argument as before, the term in the denominator is non-negative and therefore for any $\psi_J \geq 0$ there exists a $q \geq 0$ solving the equation. Therefore, $(x(q)/q)$ must converge to zero. Indeed, from the equation we can see that for $\psi_J = 0$, q must be infinite.

Convergence in the "American" System Consider average costs $C(x,q)/q$. Under the "American" system, we have that

$$\frac{C(x,q)}{q} = \frac{\theta\frac{x}{q}}{1 - exp(-\frac{rx}{q})} + \frac{\frac{f}{q} + \frac{m}{q}}{1 - exp(-\frac{rx}{q})}.$$

We want to show the limit of this expression goes to a positive number as $q \to \infty$. For the second term we have that

$$\lim_{q\to\infty}\frac{(f+m)\frac{x^*(q)}{q}\frac{1}{x^*(q)}}{1 - exp(-r\frac{x^*(q)}{q})} = \lim_{q\to\infty}\frac{(f+m)\frac{x^*(q)}{q}}{1 - exp(-r\frac{x^*(q)}{q})}\cdot\lim_{q\to\infty}\frac{1}{x^*(q)} = \lim_{\psi_A\to 0}\frac{(f+m)\psi_A}{1 - exp(-r\psi_A)}\cdot 0 = \frac{f+m}{r}\cdot 0,$$

by the multiplication rule of limits, where the first term converges to $(f+m)/r$ by L'Hopital's rule since $\psi_A \to 0$ as $q \to \infty$, and the second term converges to zero because $x^*(q) \to \infty$ as $q \to \infty$. Therefore, the overall term converges to 0.

For the first term we have that

$$\lim_{q\to\infty}\frac{\theta\frac{x}{q}}{1 - exp(-\frac{rx}{q})} = \lim_{\psi_A\to 0}\frac{\theta\psi_A}{1 - exp(-r\psi_A)} = \frac{\theta}{r},$$

where we again applied L'Hopital's rule. Therefore, overall, the average cost function under the "American" system converges to (θ/r), which is positive.

Convergence in the "Japanese" System Next consider the "Japanese" system. We have that average costs are

$$\frac{C(x,q)}{q} = \frac{\theta e^{(r+\rho)(x/q)} \frac{x}{q}}{1 - exp(-\frac{rx}{q})} + \frac{\frac{f}{q}}{1 - exp(-\frac{rx}{q})}.$$

The second term converges to zero by the same argument as before. For the first term we find

$$\lim_{\psi_J \to 0} \frac{\theta e^{(r+\rho)\psi_J} \psi_J}{1 - exp(-r\psi_J)} = \lim_{\psi_J \to 0} e^{(r+\rho)\psi_J} \cdot \lim_{\psi_J \to 0} \frac{\theta \psi_J}{1 - exp(-r\psi_J)} = 1 \cdot \frac{\theta}{r},$$

and hence average costs under the "Japanese" system asymptote to exactly the same positive limit as under the "American" system.

B Data Refinement

We refine the raw US trade transactions data from the LFTTD as follows. First, we drop all transactions that are warehouse entries, so that our dataset represents imports for consumption.[50] Second, we remove all transactions that do not include an importer identifier, an exporter identifier, an HS code, a value, a quantity, or a valid transaction date. We also drop observations with invalid exporter identifiers, e.g., those that do not begin with a letter (identifiers should start with the country name). Third, given our focus on arm's-length trade, we exclude from our analysis all related-party transactions.[51] We choose a conservative approach and exclude all relationships in which the two parties ever report being related, and all observations for which the related-party identifier is missing. Fourth, we use the concordance developed by Pierce and Schott (2012) to create time-consistent HS codes so that purchases of goods can be tracked over time. Fifth, we deflate transaction values using the quarterly GDP deflator of the Bureau of Economic Analysis, so that all values are in 2009 real dollars.[52] Sixth, since shipments of the same product between the same buyer and seller spread over multiple containers are recorded as separate transactions, we aggregate the dataset to the weekly level. Thus, multiple transactions in the same week will be summed to one. We perform this aggregation to ensure that each observation in our data reflects a genuinely new transaction rather than being part of a larger shipment. Finally, we remove outliers in terms of unit values. We compute the distribution of unit values in each HS10 by country by mode of transportation by quarter cell and drop transactions with a unit value below the 1st or above the 99th percentile.

C Construction of the Variables

C.1 Classification Regressions

In this section, we provide more details on the variables used in the classification regressions in Section 4. As discussed in the main text, we collapse all transactions of the same importer (m) - product (h) - country (c) - mode of transportation (z) quadruple in the same week into one. Therefore, a "transaction" (t) refers to a week in which the quadruple imports.

Quantity per Shipment (QPS_{mhcz}). This variable is constructed as

$$QPS_{mhcz} = \frac{\sum_t Quantity_{mhczt}}{Ntrans_{mhcz}}, \tag{A.6}$$

where $Quantity_{mhczt}$ is the quantity imported by quadruple $mhcz$ at transaction t and $Ntrans_{mhcz}$ is the total number of transactions by the quadruple.

[50] The term "Imports for consumption" denotes the total value of merchandise that physically clears customs or is withdrawn from customs bonded warehouses or U.S. Foreign Trade Zones; it does not refer to whether imports are purchased by consumers or firms.

[51] The Census Bureau defines parties as related if either party owns, controls or holds voting power equivalent to 6 percent of the outstanding voting stock or shares of the other organization.

[52] https://fred.stlouisfed.org/series/GDPDEF

Weeks between Shipments (WBS_{mhcz}). We construct this variable as

$$WBS_{mhcz} = \frac{end_{mhcz} - beg_{mhcz}}{Ntrans_{mhcz} - 1}, \tag{A.7}$$

where end_{mhcz} is the number of the week of the last transaction of the quadruple (see below for the construction of this variable), beg_{mhcz} is the number of the week of the first transaction of the quadruple, and $Ntrans_{mhcz}$ is the total number of transactions by the quadruple. The denominator represents the number of time gaps between subsequent transactions of the quadruple, which is one less than the number of transactions. If $Ntrans_{mhcz} = 1$, the average time gap cannot be computed.

Unit Value (UV_{mhcz}). We construct this variable as

$$UV_{mhcz} = \sum_t \frac{Value_{mhczt}}{Quantity_{mhczt}}, \tag{A.8}$$

where $Value_{mhczt}$ is the value imported by quadruple $mhcz$ at transaction t and $Quantity_{mhczt}$ is the corresponding quantity.

Quantity per Week (QPW_{mhcz}). This variable is constructed as

$$QPW_{mhcz} = \frac{\sum_t Quantity_{mhczt}}{end_{mhcz} - beg_{mhcz}}, \tag{A.9}$$

where $Quantity_{mhczt}$ is the quantity imported by quadruple $mhcz$ at transaction t, end_{mhcz} is the number of the week of the last transaction of the quadruple (see below for the construction of this variable), and beg_{mhcz} is the number of the week of the first transaction of the quadruple. In contrast to QPS_{mhcz}, this variable does not divide by the number of transactions but by the "flow" of imports in an average week. We note that since we require at least 5 transactions in our baseline, the beginning and end week are never the same and therefore the expression is finite.

First week (beg_{mhcz}). This variable is constructed as the first week in which a $mhcz$ quadruple transacts,

$$beg_{mhcz} = \min_t \{Week_{mhczt}\}, \tag{A.10}$$

where $Week_{mhczt}$ is the week number of the transaction, relative to the first week of 1960. Thus, for example the first week of 2016 has week number 2912.

Last week (end_{mhcz}). This variable is constructed as the last week in which a $mhcz$ quadruple transacts,

$$end_{mhcz} = \max_t \{Week_{mhczt}\}, \tag{A.11}$$

where $Week_{mhczt}$ is the week number of the transaction, relative to the first week of 1960. Thus, for example the first week of 2016 has week number 2912.

Quartile of the SPS distribution ($1\{SPS_{mhcz} = Q_n\}$). We construct this variable by taking the distribution of SPS_{mhcz} across hcz bins and assign each buyer quadruple to its respective quartile. The variable is an indicator that is equal to one if SPS_{mhcz} is in quartile n.

Average relationship length ($length_{mhcz}$). This variable is constructed in two steps. First, for every transaction t of an importer (m) - exporter (x) - product (h) - country (c) - mode of transportation (z) quintuple $mxhcz$, we compute the number of weeks past since the first transaction of any good between importer and exporter:

$$length_{mxhczt} = Week_{mxhczt} - \min_t\{Week_{mxt}\}, \tag{A.12}$$

where $Week_{mxhczt}$ is the week number of transaction t of quintuple $mxhcz$, relative to the first week of 1960. Thus, for example the first week of 2016 has week number 2912. The variable $Week_{mxt}$ is the week number of a transaction t of the buyer-seller pair mx in any good or mode of transportation.

We then compute the average relationship length of a $mhcz$ quadruple as the average of these lengths across all exporters and transactions:

$$length_{mhcz} = \frac{\sum_x \sum_t length_{mxhczt}}{Ntrans_{mhcz}}, \tag{A.13}$$

where $Ntrans_{mhcz}$ is the total number of transactions of the quadruple across all exporters.

Value per Shipment (VPS_{mhcz}). This variable is constructed as

$$VPS_{mhcz} = \frac{\sum_t Value_{mhczt}}{Ntrans_{mhcz}}, \tag{A.14}$$

where $Value_{mhczt}$ is the value imported by quadruple $mhcz$ at transaction t and $Ntrans_{mhcz}$ is the total number of transactions by the quadruple.

Dummy for differentiated goods ($Diff_h$). This variable is a dummy that is equal to one if the 6-digit HS code associated with product h is differentiated or has a reference price based on Rauch (1999) under the liberal classification.

Value per Week (VPW_{mhcz}). This variable is constructed as

$$VPW_{mhcz} = \frac{\sum_t Value_{mhczt}}{end_{mhcz} - beg_{mhcz}}, \tag{A.15}$$

where $Value_{mhczt}$ is the value imported by quadruple $mhcz$ at transaction t, end_{mhcz} is the number of the week of the last transaction of the quadruple (see above for the construction of this variable), and beg_{mhcz} is the number of the week of the first transaction of the quadruple. In contrast to VPS_{mhcz}, this variable does not divide by the number of transactions but by the "flow" of imports in an average week. We note that since we require at least 5 transactions in our baseline, the beginning and end week are never the same and therefore the expression is finite.

C.2 PNTR Regressions

In this section, we provide more details on the variables used in the baseline PNTR regressions in Section 5. As discussed in the main text, we collapse all transactions of the same importer (m) - exporter (x) - product (h) - country (c) - mode of transportation (z) quintuple in the same week into one. Therefore, a "transaction" (t) refers to a week in which the quintuple imports.

Quantity per Shipment (QPS_{mxhczp}). This variable is constructed as

$$QPS_{mxhczp} = \frac{\sum_t Quantity_{mxhczpt}}{Ntrans_{mxhczp}}, \tag{A.16}$$

where $Quantity_{mxhczpt}$ is the quantity imported by quintuple $mxhcz$ in period p (either 1995-2000 or 2002-2007) at transaction t and $Ntrans_{mxhczp}$ is the total number of transactions by the quintuple in period p.

Weeks between Shipments (WBS_{mxhczp}). We construct this variable as

$$WBS_{mxhczp} = \frac{end_{mxhczp} - beg_{mxhczp}}{Ntrans_{mxhczp} - 1}, \tag{A.17}$$

where end_{mxhczp} is the number of the week of the last transaction of the quintuple in period p (either 1995-2000 or 2002-2007) (see below for the construction of this variable), beg_{mxhczp} is the number of the week of the first transaction of the quintuple, and $Ntrans_{mxhczp}$ is the total number of transactions by the quintuple in period p. The denominator represents the number of time gaps between subsequent transactions of the quintuple, which is one less than the number of transactions. If $Ntrans_{mxhczp} = 1$, the average time gap cannot be computed. The PNTR regressions therefore require for each quintuple at least two transactions in each period p.

Unit Value (UV_{mxhczp}). We construct this variable as

$$UV_{mxhczp} = \sum_t \frac{Value_{mxhczpt}}{Quantity_{mxhczpt}}, \tag{A.18}$$

where $Value_{mxhczpt}$ is the value imported by quintuple $mxhczp$ at transaction t in period p (either 1995-2000 or 2002-2007) and $Quantity_{mxhczpt}$ is the corresponding quantity.

Quantity per Week (QPW_{mxhczp}). This variable is constructed as

$$QPW_{mxhczp} = \frac{\sum_t Quantity_{mxhczpt}}{end_{mxhczp} - beg_{mxhczp}}, \tag{A.19}$$

where $Quantity_{mhczpt}$ is the quantity imported by quintuple $mxhcz$ at transaction t in period p (either 1995-2000 or 2002-2007), end_{mxhczp} is the number of the week of the last transaction of the quintuple in period p, and beg_{mxhczp} is the number of the week of the first transaction of the quintuple in period p. In contrast to QPS_{mxhczp}, this variable does not divide by the number of transactions but by the "flow" of imports in an average week. As described above for WBS_{mxhczp}, we require for each quintuple at least two transactions in each period p so that this variable can be computed. The week number is relative to the first week of 1960. Thus, for example the first week of 2016 has week number 2912.

D Additional A vs J Classification Regressions

D.1 A versus J Classification for Thicker Relationships

Our baseline regressions in Section 4.1 are restricted to quadruples with at least five transactions. One concern might be that for quadruples that trade only few times, our variable suppliers per shipment is mismeasured because we did not observe a sufficient number of transactions. In Table A.1, we show that these results are robust to restricting observations to quadruples with at least 10 transactions.

Table A.1: A vs J Classification Regression With Buyer Quadruples With At Least 10 Transactions

Dep. var.	(1)	(2)	(3)	(4)
	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$	$\log(Length_{mhcz})$
$\log(SPS_{mh})$	0.285^{***}	0.311^{***}	-0.063^{***}	-0.489^{***}
	0.017	0.018	0.020	0.012
$\log(QPW_{mhcz})$	0.668^{***}	-0.343^{***}	-0.274^{***}	-0.083^{***}
	0.009	0.009	0.009	0.005
Observations	$2,966,000$	$2,966,000$	$2,966,000$	$2,966,000$
R-squared	0.940	0.631	0.844	0.391
Fixed effects	hcz	hcz	hcz	hcz
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on sellers per shipment defined for broader mh bins (SPS_{mh}) and total quantity shipped per week (QPW_{mhcz}). (QPS_{mhcz}), (WBS_{mhcz}), and (P_{mhcz}) are average quantity per shipment, average weeks between shipment, and average unit value. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country (c) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

D.2 A vs J Classification with More Aggregated Suppliers per Shipment

Our baseline regressions for examining whether procurement patterns are related to sellers per shipment in Section 4.1 compute SPS at the level of buyer quadruples ($mhcz$). One concern with this definition might be that buyers obtain shipments across multiple modes of transportation, and therefore procurement systems – and hence SPS – should be better defined at the mhc level. Analogously, SPS could be defined at at an even more aggregated mh level.In Tables A.2 and A.3 we show that our results relating suppliers per shipment and procurement patterns are robust to defining SPS at these higher levels of aggregation.

Table A.2: A vs J Classification Regression With SPS at mhc Level

Dep. var.	(1)	(2)	(3)	(4)
	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$	$\log(Length_{mhcz})$
$\log(SPS_{mhc})$	0.346^{***}	0.376^{***}	-0.083^{***}	-0.588^{***}
	0.011	0.011	0.015	0.012
$\log(QPW_{mhcz})$	0.687^{***}	-0.322^{***}	-0.279^{***}	-0.115^{***}
	0.010	0.010	0.009	0.003
Observations	2,966,000	2,966,000	2,966,000	2,966,000
R-squared	0.944	0.654	0.844	0.458
Fixed effects	hcz	hcz	hcz	hcz
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on sellers per shipment defined for broader mhc bins (SPS_{mhc}) and total quantity shipped per week (QPW_{mhcz}). (QPS_{mhcz}), (WBS_{mhcz}), and (P_{mhcz}) are average quantity per shipment, average weeks between shipment, and average unit value. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country (c) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.3: A vs J Classification Regression With SPS at mh Level

Dep. var.	(1)	(2)	(3)	(4)
	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$	$\log(Length_{mhcz})$
$\log(SPS_{mh})$	0.285^{***}	0.311^{***}	-0.063^{***}	-0.489^{***}
	0.017	0.018	0.020	0.012
$\log(QPW_{mhcz})$	0.668^{***}	-0.343^{***}	-0.274^{***}	-0.083^{***}
	0.009	0.009	0.009	0.005
Observations	2,966,000	2,966,000	2,966,000	2,966,000
R-squared	0.940	0.631	0.844	0.391
Fixed effects	hcz	hcz	hcz	hcz
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on sellers per shipment defined for broader mh bins (SPS_{mh}) and total quantity shipped per week (QPW_{mhcz}). (QPS_{mhcz}), (WBS_{mhcz}), and (P_{mhcz}) are average quantity per shipment, average weeks between shipment, and average unit value. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country (c) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

D.3 A versus J Classification for Different Modes of Transportation

In this section we examine whether the relationship between buyer quadruples' sellers per shipment and their procurement patterns vary by mode of transport. One concern with the results in the paper might be that the relationship between suppliers per shipment and procurement patterns holds for only some modes of transportation, and in particular not for air shipments. Results are reported in Table A.4, with the top panel focusing on vessel and the bottom on air. As indicated in the table, we find similar results for both forms of transport.

Table A.4: A vs J Classification Regression Across Mode of Transport

Dep. var.	(1)	(2)	(3)	(4)
	$\log(QPS_{mhc})$	$\log(WBS_{mhc})$	$\log(UV_{mhc})$	$\log(Length_{mhc})$
	Vessel			
$\log(SPS_{mhcz})$	0.4185***	0.4512***	-0.1722***	-0.5855***
	0.0131	0.0128	0.0094	0.0142
$\log(Quantity/Week_{mhcz})$	0.6608***	-0.3474***	-0.2631***	-0.1411***
	0.0077	0.0077	0.0089	0.0041
Observations	1,506,000	1,506,000	1,506,000	1,506,000
R-squared	0.9236	0.6855	0.8294	0.4632
	Air			
$\log(SPS_{mhcz})$	0.4104***	0.4434***	-0.0579**	-0.6528***
	0.0185	0.0195	0.0224	0.0087
$\log(Quantity/Week_{mhcz})$	0.7372***	-0.2723***	-0.3001***	-0.0830***
	0.0105	0.0108	0.0155	0.0037
Observations	1,029,000	1,029,000	1,029,000	1,029,000
R-squared	0.9328	0.6348	0.7641	0.4308

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week. (QPS_{mhcz}), (WBS_{mhcz}), (P_{mhcz}), and ($length_{mhcz}$) are average quantity per shipment, average weeks between shipment, average unit value (i.e. value divided by quantity, and average relationship length. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

D.4 A vs J Within Sellers

In this sub-section, we examine whether $mhcz$ buyer quadruples' sellers per shipment, SPS_{mhcz}, predicts theory-consistent procurement patterns *within* each of their exporter relationships. In principle, a buyer quadruple could appear "Japanese" in aggregate even if it were not with respect to each of its sellers. For example, a buyer quadruple might obtain frequent shipments from a few sellers, thus appearing to be "Japanese", but shipments within each seller might be dispersed if the buyer alternates among them. We use the following $mxhcz$-level OLS regression,

$$Y_{mxhcz} = \beta_0 + \beta_1 SPS_{mhcz} + \beta_2 \ln(QPW_{mxhcz}) + \beta_3 beg_{mxhcz} + \beta_4 end_{mxhcz} + \lambda_{xhcz} + \epsilon_{mxhcz}. \tag{A.20}$$

In this specification, Y_{mxhcz} represents procurement attributes at the buyer-seller relationship ($mxhcz$)

level, and the right-hand-side variables are defined at this level, as well, with the exception of SPS_{mhcz} which continues to be at the $mhcz$ level. We also include exporter by product by country by mode fixed effects (λ_{xhcz}) to compare buyer procurement patterns within sellers who may be heterogeneous in a number attributes, including production costs. Standard errors are clustered at the c level.

Results, reported in Table A.5, are similar to those in the previous section, providing further support for Proposition 2.3 and 6.1, as well as the use of SPS_{mhcz}. Across U.S. buyer quadruples within foreign exporters, we find that a one standard deviation decrease in SPS_{mhcz} is associated with QPS_{mxhcz} and WBS_{mxhcz} declines of 0.02 and 0.17 log points. For UV_{mxhcz}, this decline coincides with an increase of 0.01 log points. Furthermore, an increase in the total procurement quantity raises shipment size and frequency, but lowers unit import values.

Table A.5: A vs J Classification Regression Across $mxhcz$ Quintuples

Dep. var.	(1)	(2)	(3)
	$\log(QPS_{mxhcz})$	$\log(WBS_{mxhcz})$	$\log(UV_{mxhcz})$
$\log(SPS_{mhcz})$	0.100^{***}	0.696^{***}	-0.062^{***}
	0.014	0.040	0.004
$\log(QPW_{mxhcz})$	0.511^{***}	-0.171^{***}	-0.130^{***}
	0.007	0.009	0.007
Observations	$4,783,000$	$4,783,000$	$4,783,000$
R-squared	0.966	0.621	0.953
Fixed effects	$xhcz$	$xhcz$	$xhcz$
Controls	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by foreign exporter by product by country by mode of transport ($mxhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mxhcz}). (QPS_{mxhcz}), (WBS_{mxhcz}), and (P_{mxhcz}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity). All regressions include exporter by product by country by mode of transport ($xhcz$) fixed effects, control for the beginning and end week of the quintuple, and exclude buyer quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country (c) bins are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

D.5 Differentiated Products Versus Commodities

In this section we examine whether buyers are more likely to use "Japanese" procurement for differentiated goods. If these products have higher inspection costs, then by Proposition 2.1 buyers are more likely to use "Japanese" procurement for them, which implies smaller shipment size, greater frequency, and higher unit import values than products sourced under the "American" system (Proposition 2.3). Moreover, as discussed immediately above, this "Japanese" sourcing of differentiated products should be associated with fewer suppliers and longer relationships. We examine these features of the model using the commonly cited measure of product-differentiation from Rauch (1999) in the following $mhcz$-level OLS specification,

$$\overline{Y}_{mhcz} = \beta_0 + \beta_1 Diff_h + \beta_2 \ln(VPW_{mhcz}) + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{cz} + \epsilon_{mhcz}. \tag{A.21}$$

We consider four dependent variables. The first is the average number of weeks between shipments WBS_{mhcz} introduced above. We do not consider quantity per shipment or unit value here since the regression compares shipping systems across products, which are recorded in different units.[53] Instead, we use as our second dependent variable the average transaction value per shipment, VPS_{mhcz}, as a measure of average transaction size. The third variable is the average relationship length, $length_{mhcz}$. Finally, the fourth variable is a measure of the buyer's procurement type, SPS_{mhcz}. On the right-hand side, $Diff_h$ is a dummy variable indicating that product h is either differentiated or has a reference price, as opposed to being a commodity, according to the product categorization scheme proposed by Rauch (1999).[54] Because the right-hand-side variable of interest varies only at the product level, we are unable to include product fixed effects, so comparisons are made within country-mode bins by including fixed effects at that level (λ_{cz}). Since we cannot standardize quantities to be consistent across products, we control for potential scale effects using value per week (VPW_{mhcz}), rather than quantity per week, which was used above. As before, the sample period is 1992 to 2016, we include only buyer quadruples with at least 5 transactions, and standard errors are clustered at the country level.

Results, reported in Table A.6, are consistent with the model's predictions regarding inspection costs, while providing further support for the use of sellers per shipment to identify buyer types. As indicated in the first three columns of the table, we find that differentiated products are more "Japanese": they are shipped with fewer weeks between shipments, the average transaction size is smaller, and the relationships in which they are traded tend to last longer. Results in the final column provides further support for this view, as buyer quadruples encompassing differentiated goods tend to have lower sellers per shipment.

[53] For example, we cannot really compare the price of one barrel of oil to the price of one shoe.

[54] Rauch (1999) provides both a liberal and a conservative definition of differentiated goods. We use the liberal definition for the results reported in the main text, but note that these results are robust to using the conservative definition.

Table A.6: A vs J Classification Regression for Differentiated Goods

Dep. var.	(1)	(2)	(3)	(4)
	$\log(WBS_{mhcz})$	$\log(VPS_{mhcz})$	$\log(length_{mhcz})$	$\log(SPS_{mhcz})$
$Diff_h$	-0.234***	-0.227***	0.072*	-0.081**
	0.071	0.067	0.040	0.040
$\log(VPW_{mhcz})$	-0.463***	0.554***	-0.047***	-0.202***
	0.008	0.008	0.004	0.007
Observations	$2,589,000$	$2,589,000$	$2,589,000$	$2,589,000$
R-squared	0.613	0.728	0.193	0.276
Fixed effects	cz	cz	cz	cz
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by product by country by mode of transport ($mhcz$) bins on a dummy for whether the bin's product code is differentiated or reference priced according to the liberal classification by Rauch (1999) and on value shipped per week (VPW_{mhcz}). (WBS_{mhcz}), (VPS_{mhcz}), ($length_{mhcz}$), and SPS_{mhcz}) are average weeks between shipment, average value per shipment, average relationship length, and sellers per shipment. All regressions include country by mode of transport (cz) fixed effects, control for the beginning and end week of the quadruplet, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

E Additional DID Regressions

E.1 Alternate Time Periods

In the baseline DID regression results presented in the main text we compare relationships from 2002 to 2007 to those from 1995 to 2000. In this section, we show that these results also hold if we use a different post-PNTR period from 2004-2009. Table A.7 presents the results from the continuing relationship regression (11) using these alternative time periods. Table A.8 shows the results for the regression with only new relationships. Overall, we find similar results for the alternative time periods as in the main text.

Table A.7: Baseline Within-Exporter PNTR Regression: 2004-2009 vs 1995-2000

Dep. var.	(1)	(2)	(3)
	$\ln(QPS_{mxhcz})$	$\ln(WBS_{mxhcz})$	$\ln(UV_{mxhcz})$
$Post_p * China_c * NTRGap_p$	-0.1986***	-0.1986***	0.1492***
	0.027	0.027	0.024
$ln(QPW_{mxhcz})$	0.4033***	-0.5967***	-0.1333***
	0.0073	0.0073	0.0121
Observations	221,000	221,000	221,000
R-squared	0.9797	0.8732	0.9818
Fixed effects	$mxhcz, t$	$mxhcz, t$	$mxhcz, t$
Controls	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by exporter by product by country by mode of transport ($mxhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre- and post periods are 1995 to 2000 and 2004 to 2009. (QPS_{mxhcz}), (WBS_{mxhcz}), and (UV_{mxhcz}) are average quantity per shipment, average weeks between shipments, and average unit value (i.e. value divided by quantity). All regressions include $mxhcz$ and period p fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.8: New Relationship PNTR Regression: 2004-2009 vs 1995-2000

Dep. var.	(1)	(2)	(3)
	$\ln(QPS_{mxhcz})$	$\ln(WBS_{mxhcz})$	$\ln(UV_{mxhcz})$
$Post_p * China_c * NTRGap_p$	-0.0865**	-0.0865**	0.0748*
	0.0364	0.0364	0.0391
$ln(QPW_{mxhcz})$	0.4135***	-0.5865***	-0.1267***
	0.0108	0.0108	0.0151
Observations	3,158,000	3,158,000	3,158,000
R-squared	0.9675	0.8363	0.9727
Fixed effects	$mhcz, x, t$	$mhcz, x, t$	$mhcz, x, t$
Controls	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by exporter by product by country by mode of transport ($mxhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre- and post periods are 1995 to 2000 and 2004 to 2009. (QPS_{mxhcz}), (WBS_{mxhcz}), and (UV_{mxhcz}) are average quantity per shipment, average weeks between shipments, and average unit value (i.e. value divided by quantity). All regressions include $mxhcz$ and period p fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

E.2 Response to PNTR Among Continuing $mhcz$ Quadruples

In this section we examine whether the "American" to "Japanese" shift found in the last two sections persists at the more aggregate $mhcz$ quadruple level. While the preceding sections offer a tight examination of the implications of PNTR for individual buyer-seller relationships, they exclude, by design, some of the extensive margin effects of the policy change on buyers. Here, we allow for an additional margin of extensive margin adjustment, namely the formation of relationships with new suppliers that did not sell to the United States prior to PNTR. We use the following *mhczp*-level DID regression,

$$\begin{aligned} ln(Y_{mhczp}) =& \beta_1 1\{p = Post\} * 1\{c = China\} * NTR\,Gap_h + \beta_2 ln(QPW)_{mhczp} + \\ & \beta_3 \chi_{mhczp} + \lambda_{mhcz} + \lambda_p + \epsilon_{mhczp}. \end{aligned} \tag{A.22}$$

In this specification, Y_{mhczp} represents the three shipment attributes introduced above, but now computed for buyer quadruples in each period rather than buyer-seller quintuples. We include separate $mhcz$ and p fixed effects. As above, the pre-PNTR period is 1995 to 2000 and the post period is 2002 to 2007.

Results, displayed in Table A.9, show a significant decline in the average shipping size and weeks between shipments, consistent with a shift towards "Japanese" procurement. However, the increase in unit values, while positive, is statistically insignificant at conventional levels. One potential explanation for this outcome is the entry of new Chinese exporters during this period (Pierce and Schott, 2016; Amiti, Dai, Feenstra, and Romalis, 2020), including privately owned firms that tend to have lower prices than state-owned incumbents (Khandelwal, Schott, and Wei, 2013). New suppliers might also charge low, introductory prices to gain market share, further dampening unit values.

Table A.9: Within *mhcz* Quadruple PNTR DID Regression

Dep. var.	(1)	(2)	(3)
	$\ln(QPS_{mhczp})$	$\ln(WBS_{mhczp})$	$\ln(UV_{mhczp})$
$Post_p * China_c * NTR\,Gap_p$	-0.0427***	-0.0427***	0.0178
	0.0137	0.0137	0.0242
$ln(QPW_{mhczp})$	0.4361***	-0.5639***	-0.2068***
	0.0172	0.0172	0.0242
Observations	738,000	738,000	738,000
R-squared	0.9783	0.8761	0.9741
Fixed effects	$mhcz, p$	$mhcz, p$	$mhcz, p$
Controls	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by product by country by mode of transport ($mhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre-and post periods are 1995 to 2000 and 2002 to 2007. (QPS_{mhczp}), (WBS_{mhczp}), and (UV_{mhczp}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period p. All regressions include $mhcz$ and period p fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruples with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

E.3 PNTR Regression Without Quantity Control

In this sub-section, we examine whether our within-exporter PNTR results from Section 5.2 also hold when we do not control for quantity per week, QPW_{mxhczp}. One concern with our analysis could be that by conditioning on quantity, we do not take into account that PNTR also affects the quantity traded, which could in turn affect the procurement system. Our results in Table A.10 show that we still find a decline in the quantity per shipment and an increase in the unit value even without the quantity control.

Table A.10: Baseline Within-Exporter PNTR Regression Without Quantity: 2002-2007 vs 1995-2000

Dep. var.	(1)	(2)	(3)
	ln(QPS_{mxhczp})	ln(WBS_{mxhczp})	ln(UV_{mxhczp})
$Post_p * China_c * NTRGap_p$	-0.2753***	-0.0625**	0.1186***
	0.0187	0.0303	0.0174
Observations	439,000	439,000	439,000
R-squared	0.97	0.66	0.98
Fixed effects	$mxhcz, p$	$mxhcz, p$	$mxhcz, p$
Controls	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by exporter by product by country by mode of transport ($mxhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre-and post periods are 1995 to 2000 and 2002 to 2007. (QPS_{mxhczp}), (WBS_{mxhczp}), and (UV_{mxhczp}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period p. All regressions include $mxhcz$ and period p fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

E.4 PNTR Regression For All Relationships

In this sub-section, we re-run our relationship-level PNTR regression (11) using all relationships that appear in either the pre-PNTR or the post-PNTR period for all buyer quadruples and sellers that appear in both. Specifically, we run the regression using separately importer-product-country-mode of transportation ($mhcz$) fixed effects and exporter (x) fixed effects. Our results in Table A.11 indicate that PNTR leads to a decline in the quantity per shipment and an increase in the unit value for this set of relationships.

Table A.11: Within-Exporter PNTR Regression Using All Relationships: 2002-2007 vs 1995-2000

Dep. var.	(1)	(2)	(3)
	$\ln(QPS_{mxhczp})$	$\ln(WBS_{mxhczp})$	$\ln(UV_{mxhczp})$
$Post_p * China_c * NTRGap_p$	-0.1311***	-0.1311**	0.0782***
	0.0163	0.0163	0.0214
$ln(QPW_{mxhcz})$	0.4068***	-0.5932***	-0.1303***
	0.0116	0.0116	0.0162
Observations	4,023,000	4,023,000	4,023,000
R-squared	0.97	0.66	0.98
Fixed effects	$mhcz, x, p$	$mhcz, x, p$	$mhcz, x, p$
Controls	Yes	Yes	Yes

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by exporter by product by country by mode of transport ($mxhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre-and post periods are 1995 to 2000 and 2002 to 2007. (QPS_{mxhczp}), (WBS_{mxhczp}), and (UV_{mxhczp}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period p. All regressions include $mhcz$, exporter x, and period p fixed effects, control for the beginning and end week of the quadruplet as well as all variables needed to identify the DID term of interest, and exclude quadruplets with less than 5 shipments. Standard errors, adjusted for clustering by country, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

F Market Clearing Conditions

This section provides the market clearing conditions of the multi-country model described in Section 6.

Goods market clearing implies for each manufactured good ω

$$\sum_n \sum_i \sum_s I_{ni,s}(\omega) x^*_{ni,s}(\omega) = \sum_n \sum_i \sum_s I_{ni,s}(\omega) \int_0^{x^*_{ni,s}(\omega)/q_n(\omega)} q_n(\omega) dt \quad \forall \omega, \tag{A.23}$$

where the left-hand side of the equation indicates worldwide production of product ω in one order cycle and the right-hand side is worldwide consumption over the same cycle.

The market for non-manufactured goods clears as well,

$$\sum_n Z_n = \sum_n a_n L_n^{NM}. \tag{A.24}$$

Finally, labor market clearing in each country requires that

$$\begin{aligned} L_n = & \sum_{n'} \sum_s I_{n'n,s}(\omega) \int_0^1 \frac{\bar{\theta}}{\Upsilon_n(\omega)} q_{n'}(\omega) d\omega + f_n \sum_{n'} \sum_s I_{n'n,s}(\omega) \int_0^1 \frac{q_{n'}(\omega)}{x^*_{n'n,s}(\omega)} d\omega \\ & + \sum_i I_{ni,A} \int_0^1 (\omega) m(\omega) \frac{q_{n'}(\omega)}{x^*_{n'n,s}(\omega)} d\omega + L_n^{NM} \quad \forall n \in N, \end{aligned} \tag{A.25}$$

where the left-hand side is total labor supply in country n, and on the right-hand side we have labor used in manufacturing production, labor used for fixed costs, labor used for inspections, and non-manufacturing labor, respectively. Since the fixed costs and the inspection costs are paid for each shipment, we scale these costs by the number of shipments per period.

G Equilibrium Solution Algorithm

We solve for the equilibrium of the model in several steps. We discretize the product space to $\Omega = 5,000$ products for the estimation algorithm.

1. Define a four-dimensional grid with $(K_1 \times K_2 \times K_3 \times Q)$ grid points. Let $\mathbf{k} \equiv (k_1, k_2, k_3, q_k)$ denote a given grid point. Solve numerically for the average costs $AC(\mathbf{k})$ at each grid point under each system, using equation (5), i.e.

$$AC_A(\mathbf{k}) = \min_x \left(\frac{r}{q_k}\right) \frac{k_1 + k_2 x}{\left[1 - e^{-rx/q_k}\right]} \tag{A.26}$$

   and

$$AC_J(\mathbf{k}) = \min_x \left(\frac{r}{q_k}\right) \frac{k_1 + e^{(r+k_3)x/q_k} k_2 x}{\left[1 - e^{-rx/q_k}\right]}. \tag{A.27}$$

   We denote by $x_A(\mathbf{k})$ and $x_J(\mathbf{k})$ the cost-minimizing shipment sizes under each system at grid point $\mathbf{k}$.

2. Initiate the model by drawing an inspection cost $m(\omega)$ for each product ω from $G(m)$ and by drawing a productivity $\Upsilon_i(\omega)$ from $F_i(\Upsilon)$. Also set the trade war arrival rates ρ_{ni} for each country pair.

3. Compute the average cost for each origin country (i)-destination country (n)-product (ω) triplet under each system s, $AC_{ni,s}(x^*_{ni,s}, q_k)$, at each of the Q quantity grid points. Specifically, we map each $ni\omega$'s draw $(m(\omega), \Upsilon_i(\omega), \rho_{ni})$ to an estimated average cost using linear interpolation on the grid of average costs computed in Step 1, where under the "American" system we use $k_1 = f_i w_i + m(\omega) w_n$, $k_2 = \frac{\bar{\theta}}{\Upsilon_i(\omega)} w_i$ and under the "Japanese" system we use $k_1 = f_i w_i$, $k_2 = \frac{\bar{\theta}}{\Upsilon_i(\omega)} w_i$, and $k_3 = \rho_{ni}$. Similarly, obtain the shipment sizes from linear interpolation on the grid of shipment sizes computed in Step 1.

4. Determine the cost minimizing system and origin country at each quantity q_k for each destination-product market $n\omega$, using equation (15). This traces out the average cost curve of each tuple.

5. Guess an initial manufacturing price index in each destination country, $p_n(0)$, and an initial total income, $W_n(0)$.

6. Begin iteration, t=0.

(a) Compute each destination-product market $n\omega$'s demand curve, using utility maximization, by computing for each q_k

$$p_n(\omega; q_k, t) = \left(\frac{\alpha W_n(t)}{q_k}\right)^{\frac{1}{\sigma}} p_n(t)^{\frac{\sigma-1}{\sigma}}. \tag{A.28}$$

(b) Find the intersection between supply and demand curve in each market, using linear interpolation between grid points, to obtain the equilibrium $(p_n^*(\omega), q_n^*(\omega))$. If there are several intersections, find the last intersection at which the demand curve intersects the supply curve from above.

(c) Use the equilibrium prices in each market to compute a new price index, $p_n(t+1)$.

(d) Determine the labor used for production, fixed costs, and inspection costs in each period, and compute the labor used in non-manufacturing L_n^{NM} from the labor market clearing condition (A.25). Verify that this labor is non-negative in all markets.

(e) Compute the total income of each country n, which is equal to labor income $w_n L_n$ plus profits under the "Japanese" system, see equation (16).

(f) Repeat the iteration steps with the new values of $p_n(t+1)$ and $W_n(t+1)$ and iterate to convergence.

We choose to solve for the average costs on a grid for efficiency reasons. While in principle it would be possible to solve numerically for the average costs of each $ni\omega$ tuple directly, thus combining steps 1-3, in practice the numerical solution of the average cost problem is quite time consuming. Since we have to solve the model many times using our MCMC algorithm, this would lengthen the estimation time considerably. By separating out the steps, we can in a first step numerically solve the cost minimization problem for a large number of combinations of parameter values and then save the results prior to the actual estimation procedure. During the estimation we then only use linear interpolation on this estimated cost grid without doing any further numerical solution, which is much faster.

H Construction of Empirical Moments

In this section, we describe how each of the moments used in the estimation is constructed.

1. Share of Chinese imports in domestic sales

We target the U.S. import penetration from China in 2016. We obtain U.S. imports of goods from China in 2016 from the Census Bureau[55], and obtain gross output in the manufacturing sector in 2016 from the Bureau of Economic Analysis (BEA).[56] We also obtain total U.S. imports and exports

[55]https://www.census.gov/foreign-trade/balance/c5700.html.

[56]https://apps.bea.gov/iTable/iTable.cfm?reqid = 150&step = 2&isuri = 1&categories = gdpxind.

of goods in 2016 from the Census Bureau. We then construct the target

$$IP_{CN} = \frac{\text{Imports}_{CN}}{\text{Domestic output} + \text{Total imports} - \text{Total exports}}, \tag{A.29}$$

where Imports_{CN} are U.S. imports from China, Domestic output denotes gross output in the manufacturing sector, and Total imports and Total exports are total U.S. imports and exports of goods.

2. Share of rest of the world imports in domestic sales

We target the U.S. import penetration from the rest of the world in 2016. We construct this target analogously to import penetration from China. We obtain U.S. imports of goods from China in 2016 from the Census Bureau, and obtain gross output in the manufacturing sector in 2016 from the Bureau of Economic Analysis (BEA). We also obtain total U.S. imports and exports of goods in 2016 from the Census Bureau. We then construct the target

$$IP_{RoW} = \frac{\text{Imports}_{RoW}}{\text{Domestic output} + \text{Total imports} - \text{Total exports}}, \tag{A.30}$$

where Imports_{RoW} are U.S. imports from the rest of the world, Domestic output denotes gross output in the manufacturing sector, and Total imports and Total exports are total U.S. imports and exports of goods.

I Estimation Algorithm and Outcomes

The objective is to find a parameter vector ϕ^* that solves

$$\arg\min_{\phi \in \mathbb{F}} \sum_x T(m_x(\phi), \hat{m}_x) \tag{A.31}$$

where $T(\cdot)$ is the percentage difference between the model, $m_x(\phi)$, and data, $\hat{m}_x$, moments, and $\mathbb{F}$ is the set of admissible parameter vectors, which is bounded to be strictly positive and finite. In the choice of the function $T(m_x(\phi), \hat{m}_x)$ we follow Jarosch (2016) and **?** and minimize the sum of the percentage deviations between model-generated and empirical moments.

The minimization algorithm that we use to solve the problem combines the approaches of Jarosch (2016) and **?**, and **?**, adapted to our needs.

We simulate, using Markov Chain Monte Carlo for classical estimators as introduced in **?**, 200 strings of length 10,000 (+ 1,000 initial scratch periods used only to calculate posterior variances) starting from 200 different guesses for the vector of parameters ϕ_0. In the first run, we choose the initial guesses to span a large space of possible parameter vectors. In updating the parameter vector along the MCMC simulation, we pick the variance of the shocks to target an average rejection rate of 0.7, as suggested by **?**. The average parameter values across the 200 strings for the last 1,000 iterations provide a first estimate of the vector of parameters. We then repeat the same MCMC procedure, but we start each string from the parameter estimates of the first step. We pick our final estimates as the average across the parameter vectors, picked from all strings, that are associated with the 100 smallest values of the likelihood functions.

Figure XX illustrates our approach. The black dotted line shows the density function of the last 1,000 iterations across all strings. We pick the optimal parameter following **?** and select the vector of parameters that minimizes the objective function among all our draws.[57] Our estimates are shown with red dotted lines in the figure. These correspond to the estimates reported in Table 11 in the main text. Finally, the blue density functions shows the density, across all strings, of the 10 best outcomes within each string. This density provides a visual representation of the tightness of our estimates, which are, in general, quite good. It is also relevant to notice that all the densities are single-peaked, which suggests that the model is, at least locally, tightly identified.

Figure YY illustrates how each parameter is identified.

[57] More precisely, we take the average across the 100 best outcomes across all the draws.