Tariff Rate Uncertainty and the Structure of Supply Chains*

Sebastian Heise † Justin R. Pierce ‡ Georg Schaur §
Peter K. Schott ¶

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Abstract

We show that reductions in the probability of a trade war promote long-term importer-exporter relationships that ensure provision of high-quality inputs via incentive premia. Empirically, we introduce a method for distinguishing between these "Japanese" versus spot-market (i.e., "American") relationships in customs data, show that their use varies intuitively across trading partners and products, and find that "Japanese" importing from China increases after a reduction in the possibility of a trade war. Extending the standard general equilibrium trade model to encompass potential trade wars and relational contracts, we estimate that eliminating "Japanese" procurement reduces welfare about a third as much as moving to autarky. (JEL Codes: F13, F14, F15, F23)
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†Federal Reserve Bank of New York; sebastian.heise@ny.frb.org.
‡Board of Governors of the Federal Reserve System; justin.r.pierce@frb.gov
§University of Tennessee; gschaur@utk.edu
¶Yale School of Management & CEPR & NBER; peter.schott@yale.edu

1 Introduction

Since the early 1990s, the rapid expansion of global value chains has promoted a substantial increase in world trade, boosted aggregate productivity, and supported an unprecedented convergence in rich and poor country incomes (Johnson, 2018; World Bank, 2020; Antras and Chor, 2022). For much of this time, the international spread of production networks was bolstered by trade liberalization and policy stability. Rising protectionism in recent years, however, has dampened firms' enthusiasm for global supply chains, threatening the welfare gains of previous decades.[1] In this paper, we demonstrate that the welfare effect of an increase in the *probability* of trade restrictions depends on firms' use of different types of procurement systems.

We develop a model of global sourcing that builds upon the partial-equilibrium framework for domestic supply chains introduced by Taylor and Wiggins (1997). In that framework, buyers choose an optimal order pattern, payments, and inspections to procure inputs from sellers that benefit from evading quality standards. Buyers' cost-minimizing strategy is one of two systems. Under the "Japanese" system, buyers motivate a 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 from a parade of lowest-cost sellers in the spot market. Costly inspections and enforceable contracts deter cheating. Lower inspection costs favor the "American" system, while factors supporting firms' ability to form long-term relationships favor the "Japanese" system. We hereafter refer to the "Japanese" and "American" systems as J and A.

In the first part of the paper, 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 probability of a trade war. In equilibrium, each buyer procures and distributes its product using the system that minimizes costs. We show that increases in the probability of a trade war reduce the likelihood that buyers choose J procurement because it shortens the expected length of buyer-seller relationships, thereby raising the premia buyers must pay sellers to incentivize high quality under the J system.

In the second part of the paper, we document the prevalence of J sourcing among

[1] See, for example, Amiti et al. (2019), Fajgelbaum et al. (2019), Flaaen and Pierce (2019), Bown et al. (2021), and Flaaen et al. (2020).

US importers using transaction-level US import data that record both the number of foreign exporters with which US importers trade as well as the values and quantities (and therefore the unit values) associated with each shipment. Guided by our model, we classify importers as using either J or A procurement based on the number of foreign suppliers from which they purchase a particular product from each country over the 1992 to 2016 sample period. A lower ratio of suppliers to the number of shipments indicates more repeat purchases from the same seller, hence a higher likelihood of the J system. Intuitively, we find that J importing is most prevalent from Japan and Mexico, for products classified as transportation and machinery, and for imports obtained by manufacturing versus service firms. We then show, consistent with the model, that buyers with a lower ratio of suppliers to shipments do indeed receive smaller, more frequent shipments at a higher price than A buyers of the same product. J buyers also tend to be larger, pay higher wages, and have lower inventory to sales ratios. These results provide the first systematic empirical evidence of the J and A procurement patterns as highlighted by Taylor and Wiggins (1997).[2]

In the third part of the paper, we provide evidence of a switch towards J procurement among US importers and Chinese exporters after a 2001 change in US trade policy that substantially reduced the probability of a trade war between the two counties. Our triple difference-in-differences specification, which asks whether US 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 the policy (third difference), provides support for the model along two dimensions. First, we show that imports by importers of more-exposed products from China become relatively smaller, more frequent, and increase in unit value after the change in policy, consistent with a shift to J. A one standard deviation increase in exposure to the policy is associated with a relative decline in shipment quantity of 4.5 percent and a relative increase in shipment frequency and shipment unit value of 3.9 percent and 2.1 percent. Second, we find that US importers of more-exposed products exhibit a relative reduction in sellers per shipment. Both results indicate a shift from A to J procurement among the products that benefit most from the elimination of future tariff threats.

In the final part of the paper, motivated by recent events such as Brexit and US

[2]Building on our approach, Cajal-Grossi et al. (2023) find further support for the sellers per shipment measure analyzing buyer-seller relationships in Bangladesh's garment industry.

-China "de-risking", we embed our procurement framework in an Eaton and Kortum (2002) model of trade to provide the first assessment of the impact of trade policy uncertainty on relational contracting, i.e., repeated transactions between buyers and sellers under an informal agreement. In our setup, sourcing is governed by bilateral trade war arrival rates in addition to standard cross-country differences in productivity, as they affect the relative costs of procurement under the two systems. Quantitative simulations of the model reveal that an increase in the probability of a trade war that is sufficient to eliminate J-style procurement reduces US welfare about one third as much as placing the US in autarky.

Literature

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 et al., 2024; Handley and Limão, 2022) by identifying procurement systems as a new channel through which uncertainty can influence trade patterns and welfare. Our finding of a relationship between procurement system switching and unit values highlights a novel source of price variation in response to changes in trade policy that goes beyond the quality premiums and markups studied in the existing literature (Schott, 2004; Verhoogen, 2008; Khandelwal, 2010; Hallak and Schott, 2011; Kugler and Verhoogen, 2012; Antoniades, 2015; Manova and Yu, 2017). Our model also demonstrates that the distributional implications of changes in policy uncertainty depend on firms' procurement strategies. Firms choosing to enter relational contracts, hence using J sourcing, are more sensitive to increases in the probability of a trade war than firms that procure on the spot market.

Second, we contribute to greater understanding of the organization of global value chains (Antràs et al., 2017; Antràs and Chor, 2018; Antras and Chor, 2022), as well as a larger literature on incomplete contracts, imperfect contract enforcement, and information asymmetries (Antràs, 2003; Antràs, 2005; Grossman and Helpman, 2004; Spencer, 2005; Feenstra and Hanson, 2005; Antràs and Helpman, 2008; Kukharskyy and Pflüger, 2010). In contrast to much of the research in this area, we consider different types of procurement systems, J versus A, 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. Our work builds upon the literature on relational contracting as an alternative to integration (Defever et al., 2016; Kukharskyy, 2016), where the pattern of trade between buyers and sellers is usually governed by idiosyncratic time preferences.[3] In our model, by contrast, discount rates are common and firms choose between procurement systems based on inspection costs and policy stability. As a result, our model links shipment patterns to policy in a manner amenable to empirical inquiry using transaction-level trade data.

Third, our results relate to analyses of importer-exporter trade flows demonstrating that high fixed per-shipment trade costs reduce shipping frequency, thereby raising inventories in a manner that influences firms' adjustment to trade shocks (Alessandria et al., 2010; Alessandria et al., 2011; Kropf and Sauré, 2014; Hornok and Koren, 2015a; Hornok and Koren, 2015b; Békés et al., 2017). Here, we document that firms that source under the A system have higher inventories, and show that trade policy uncertainty can be an important barrier to firms' efforts to reduce inventory costs, as it promotes use of the A system. We estimate inspection costs of 0.4 percent of the transaction value for the average import transaction, about one tenth of our estimated average fixed cost per shipment.

Finally, we examine the consequences of relational contracting in general equilibrium by extending Eaton and Kortum (2002) along two dimensions. First, we have product prices depending on the probability of a trade war as well as supplier productivity. Second, we have increasing returns to scale in procurement costs due to shipment-specific logistics fees. Quantification exercises reveal that reducing buyers' ability to form J relationships due to a higher trade war probability lowers welfare by raising prices, akin to an adverse productivity shock. While we focus on changes in the probability of a trade war, our mechanism applies to any factor that undermines sellers' beliefs about the viability of long-term relationships with buyers, e.g., uncertainty about shipment arrival due to corruption, pandemics, or port disruptions.

We examine J sourcing theoretically and empirically in Sections 2 through 4. Sections 5 through 7 extend our model to general equilibrium and perform counterfactuals. Section 8 concludes. An appendix provides additional detail and results.

[3] For recent evidence on the potential of relational contracting in Rwanda's coffee and Kenya's rose industries see Macchiavello and Morjaria (2020) and Macchiavello and Morjaria (2015).

2 Extending Taylor and Wiggins, 1997

Quality control and incomplete contracts are a common problem in firms' procurement decisions. Taylor and Wiggins (1997) provide a framework that focuses on an arm's-length solution to these challenges.[4] In their theory, a buyer repeatedly seeks to obtain high-quality inputs from a supplier whose effort is unobservable.[5] Their solution to this problem is one of two optimal contracts. Under the A 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 J system, by contrast, buyers offer sellers a price premium over a long-term relationship as an incentive to deter cheating. The Taylor and Wiggins (1997) framework is particularly suitable to our context because it broadly characterizes typical procurement strategies (Helper and Sako, 1995) and, linking incentive premia to potential trade wars, allows us to examine the effect of trade policy stability on international shipping patterns and welfare.

2.1 The Procurement Problem

The Seller's Problem: There is a single country populated by a continuum of homogeneous sellers able to produce the same good.[6] To complete a production run (i.e., produce one shipment) a seller hires labor l at wage $w = 1$ to produce and deliver output $x = \frac{\Upsilon}{\theta} l$, where Υ is a seller's productivity and θ represents her product's level of quality. The unit input requirement, $\frac{\theta}{\Upsilon}$, allows for variation in quality, giving rise to a "quality control" problem.[7] Sellers choose between discrete quality levels, $\theta \in \left\{\underline{\theta}, \overline{\theta}\right\}$, where lower quality is less costly to produce. To complete the shipment, the seller absorbs f units of labor for per-shipment logistics services, including transport costs.[8] The seller's total costs for each production and delivery cycle are therefore $x\frac{\theta}{\Upsilon} + f$.

The Buyer's Problem: Homogeneous buyers with complete bargaining power procure

[4]Firm integration is another but potentially very costly means of addressing these issues (Antràs, 2003; Antràs, 2005; Antràs and Helpman, 2008). China, for example, requires foreign ventures to include a domestic partner, while the United States (and other developed countries) mandate national security reviews.

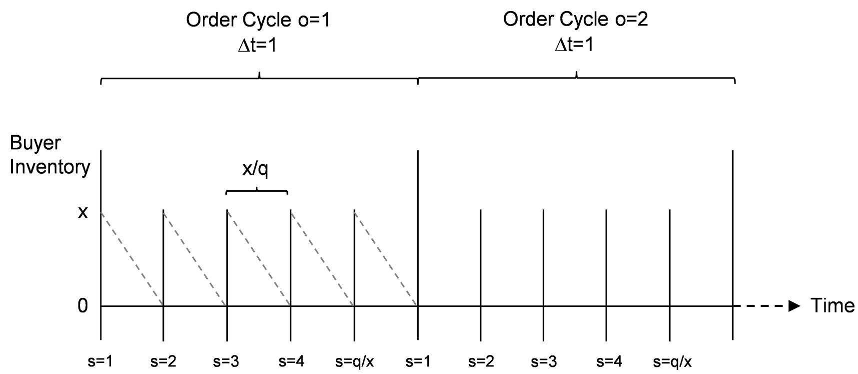
[5]This problem falls into the class of repeated games with incomplete information (Kandori, 2002).

[6]We extend the model to multiple products and sellers in multiple countries in Section 5.

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

[8]Recent evidence emphasizes per-unit and -shipment specific delivery costs (Hummels and Skiba, 2004; Martin, 2012; Kropf and Sauré, 2014; Hornok and Koren, 2015a; Hornok and Koren, 2015b).

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, arriving every x/q units of time apart.

a seller's output and distribute it to consumers. 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 5. Consequently, there are q/x shipments during each period. Figure 1 summarizes the shipment and consumption pattern. If quality is less than desirable, then no demand arrives and buyers must dispose of the obsolete shipment without recompense. Following Taylor and Wiggins (1997), the buyer seeks to ensure the desired level of quality using either an A or a J procurement system.

In the A system, buyers inspect each shipment, and inspections reveal product quality with certainty.[9] Inspection costs m_A for each shipment are fixed.[10] 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 exactly break even and participate, where

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

[9] Taylor and Wiggins (1997) allow for probabilistic inspections and derive limit theorems for small discount rates. Our simplification facilitates analytical tractability when we extend discount rates for the possibility of trade wars.

[10] "[I]t costs the same to have 20 pallets inspected as it does just one." See "What a Year of Brexit Brought UK Companies: Higher Costs and Endless Forms," New York Times, December 29, 2021.

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 $v_A(x_A, \theta) + m_A$.

J procurement motivates the production of high quality via an incentive premium and the value of a long-term relationship. This value depends upon the relationship's longevity. Let trade policy shocks that break buyer-seller relationships, e.g. tariff escalation to prohibitive levels, arrive at a constant rate, ρ.[11] Then, relationships survive over a shipment cycle with probability $e^{-\frac{\rho x}{q}}$.[12] Our focus is on trade policy but other shocks including natural disasters may have similar consequences (Boehm et al., 2019).

If $e^{-\frac{\rho x}{q}} < 1$, then firms are uncertain about whether future trade policy will sustain relationships and a greater arrival rate of trade wars, ρ, increases the separation probability.[13] Let r be the per-period interest rate and $v_J(x_J, \theta)$ be the payment the buyer sets under the J system for each shipment. With continuous compounding, the expected discounted value of the relationship is then $\frac{v_J(x_J,\bar{\theta})}{1-e^{-(r+\rho)x_S/q}}$.[14]

If the buyer does not observe product quality until the shipment is received and the payment is made, then, to guarantee desired quality, he sets a per-shipment payment such that the seller's net present value of the continued relationship exceeds the one-time profit from cheating on quality, $\frac{v_J(x_J,\bar{\theta})-f-\frac{\bar{\theta}}{\Upsilon}x_J}{1-e^{-(r+\rho)x_J/q}} \geq v_J(x_J, \bar{\theta}) - f - \frac{\underline{\theta}}{\Upsilon}x_J$. Rearranging, buyers under the J system set the per-shipment payment

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

The per-unit premium $\left[e^{(r+\rho)x_J/q} - 1\right](\bar{\theta} - \underline{\theta})\frac{1}{\Upsilon}$ incentivizes quality. A key feature of the J system is that more stable trade relationships (i.e., a lower ρ) with repeated smaller shipments, x_J, sent more frequently reduce the premium necessary to guar-

[11] In a potential trade war average tariffs are estimated at 63 percent worldwide (Ossa, 2014).

[12] Relationships thus break with probability $F(t) = 1 - e^{-\rho t}$ over interval t (Wooldridge, 2002, p. 688). At the product level, 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).

[13] Handley and Limão (2017) consider trade policy where tariffs may either go up or down. In our case, the uncertainty is w.r.t. greater tariffs that break relationships.

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

antee desired quality. Long-term relationships are optimal in the Japanese system because they increase the incentive to provide quality.

Buyers choose between the A and J system by comparing long-term expected revenues and costs taking into account that trade wars will result in a loss of profits. At a given market price p, long-term expected profits in the two procurement systems are then given by

$$\pi_s^b = \left[\int_0^{x_s/q} e^{-rt} pq\, dt - v_s(x_s, \overline{\theta}) - m_s\right] / \left[1 - e^{-(r+\rho)x_s/q}\right] \quad \text{s}\in\{\text{J,A}\} \tag{3}$$

where discounted revenues per 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 determine the optimal procurement system. In equilibrium, buyers' profits equal zero (see Section 5). Therefore, the market price must equal average costs, $AC_s\,(x_s, q)$, and employing (3) set equal to zero we obtain

$$p_s = AC_s\,(x_s, q) = \left(\frac{r}{q}\right) \frac{v_s(x_s, \bar{\theta}) + m_s}{[1 - e^{-rx_s/q}]} \quad \text{s}\in\{\text{J,A}\}. \tag{4}$$

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,

$$\frac{v_s'(x_s, \bar{\theta})}{1 - e^{-rx_s/q}} = \frac{\left[v_s(x_s, \bar{\theta}) + m_s\right] \frac{r}{q} e^{-rx_s/q}}{\left(1 - e^{-rx_s/q}\right)^2} \quad \text{s}\in\{\text{J,A}\}. \tag{5}$$

The firm optimally procures x_s^* such that the discounted value of higher costs associated with a small increase in order size (left-hand side) equals the savings from an increased discount factor due to spacing these larger orders further apart in time (right-hand side).[15]

The buyer compares average procurement costs evaluated at the optimum, $AC_s\,(x_s^*, q)$, to determine the cost-minimizing procurement system. If $\bar{\theta} - \underline{\theta} = 0$ and with $m_A = 0$, then there is no incentive problem and costs in both systems are identical. Com-

[15] Supplemental Appendix J.1 shows that an interior solution to the first order condition is a unique cost minimizer for $0 < rx/q < 1$. The Supplemental Appendix is available on the authors' websites. It is not for publication and provides additional results not central for the argument.

pared to this benchmark case, differentiating equation (4) under the J system with respect to $\underline{\theta}$ and ρ using the envelope theorem shows that average procurement costs in the J 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 $\frac{\partial AC_J\left(x_J^*, q\right)}{\partial \rho} \geq 0$.[16] In the A system, differentiating (4) with respect to m shows that average costs increase with inspection costs m. Importantly, as $m \to \infty$, we have $AC_A\left(x_A^*, q\right) \to \infty$ because average costs grow without bound, $\frac{\partial AC_A\left(x_A^*, q\right)}{\partial m} = \frac{1}{1-e^{-\frac{rx_A^*}{q}}} > 1$. We obtain 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 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 J system is chosen for $m > m^*$.*

Proof. See Appendix A.2. □

This proposition highlights that the arrival rate of trade wars affects the average procurement cost under the J system and buyers' endogenous choice of the procurement system. Starting at a level of inspection cost m slightly below m^*, a reduction in ρ lowers average costs under the J system and reduces the threshold inspection cost m^* at which procurement costs under both system are the same, causing the buyer to switch from the A to the J system if m^* falls below m.[17]

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. 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 bounded, 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 J 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 A system.*

[16] See Appendix Section A.1 for the proof.

[17] Existing theories of relational contracts in trade rely on exogenous heterogeneity in discount rates to determine relationship-based transactions (Kamal and Tang, 2015; Defever et al., 2016; Kukharskyy, 2016). In our framework, buyers endogenously determine the effective discount rate of rx_s/q by choosing the optimal procurement system and order size in response to inspection costs and the probability of a future trade conflict.

Proof. See Appendix A.3. □

Under the J procurement system an increase in ρ raises the incentive premium. As a result, variable procurement costs increase and buyers re-optimize by lowering shipment sizes (i.e., raising shipping frequency). Unit values increase because fixed per-shipment costs are spread over smaller shipment sizes. Instead, an increase in the inspection cost m raises fixed per-shipment costs under the A system, and buyers re-optimize by increasing per-shipment quantities (i.e., decreasing shipping frequency). The unit value paid to the seller must decrease in the A system since the fixed cost f is spread over more units.

We can now rank shipping frequencies and unit values across the two systems. If $\bar{\theta} - \underline{\theta} = 0$ and $m_A = 0$, then the A and J procurement systems are identical. An increase in $\bar{\theta} - \underline{\theta}$ raises variable shipment costs under the J system, leading buyers to increase their shipping frequency by lowering the shipment size. Unit values increase because fixed costs are spread over fewer units. Under the A system, Proposition 2.2 shows that an increase in inspection costs raises the shipment size, and hence shipping frequency and unit values decrease. Therefore, if $\bar{\theta} - \underline{\theta} > 0$ and $m \geq 0$, then shipping sizes are greater in the A system and unit values are greater in the J system. This reasoning forms the basis of our third proposition.

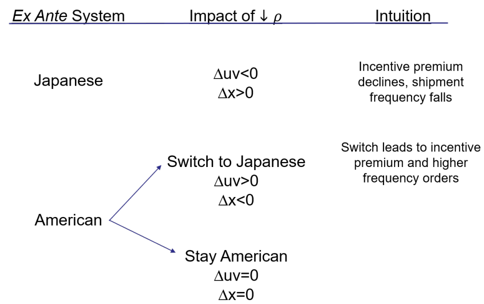
Proposition 2.3. *Batch sizes in the A system are greater than in the J system, $x_A^* > x_J^*$, and therefore time between shipments is greater under the A system, $x_A^*/q > x_J^*/q$. Unit values in the J system are greater than in the A system, $v_J(x_J, \bar{\theta})/x_J > v_A(x_A, \bar{\theta})/x_A$.*

Proof. See Appendix A.4. □

Figure 2 illustrates the predictions of a lower likelihood of trade war (a decrease in ρ) according to Proposition 2.2 and 2.3. The effect depends on whether the adjustment takes place within the J system or via a switch from the A to the J system. Within the J system, unit values fall, shipment sizes increase, and shipping frequency declines. Within the A system we expect no impact on prices, quantities, or frequencies. If a lower trade war arrival rate triggers a switch from A to J procurement, then we predict a decrease in shipment sizes and an increase in the unit value.

In Section 3, we show that the frequency of US importing from China under the J system is relatively low in the first part of our sample period, but that it rises

Figure 2: Impact of A Decline in the Probability of Trade Conflict (ρ)



Notes: Figure illustrates the impact of a change in the arrival rate of a trade war, ρ on shipment unit values (uv) and quantities (x) under both systems where, e.g., $\Delta uv < 0$ indicates a decline in unit value.

over time. Consistent with this finding, Section 4 shows that a plausibly exogenous reduction in ρ *vis à vis* China primarily results in A to J switching.

3 Data on J Importers

We use the US Census Bureau's Longitudinal Foreign Trade Transaction Database (LFTTD) to identify J exporters and examine the predictions of the model introduced above. Our dataset tracks every US import transaction from 1992 to 2016 and includes: the dates the shipment left the exporting country and arrived in the United States; identifiers for the US 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.[18] We perform standard data cleaning and use the concordance developed by Pierce and Schott (2012) to create time-consistent HS codes. Given our focus on arm's-length trade, we drop all related-party transactions. 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. For more detail on our data preparation, see Appendix Section B.

Our analysis below focuses on "buyer quadruples" that group shipments of a ten-digit HS product (h) imported by a US importer (m) from origin country (c) shipped

[18]We focus on vessel, rail, road, and air, dropping the small fraction of transactions that are transported by other means, e.g., hand-carried by passengers. See Bernard et al. (2009) for further information on the LFTTD and Kamal and Monarch (2018) for more detail on the foreign firm identifier.

via mode of transportation (z).[19] Since our theory requires that we observe repeated shipments to learn about the procurement system, we exclude buyer quadruples with fewer than five shipments in our analysis.[20] Our sample represents more than 80 percent of all arm's length trade and contains almost 3 million $mhcz$ quadruples between 1992 and 2016. There are nearly 22 million "buyer-seller relationships" associated with these bins, i.e., the number of $mxhcz$ quintuples, where x denotes the exporter. Table A.1 in Appendix B provides an overview.[21]

Table 1 summarizes the $mhcz$ quadruples, which are the focus of our study in the next section. The first four rows of the table reveal that from 1992 to 2016, the average $mhcz$ bin traded 1.9 million dollars (in 2009 units), 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 across the relationships within a quadruple ($length_{mhcz}$) averaged 36 thousand dollars, 24 weeks and 181 weeks, respectively.[22]

3.1 Sellers per Shipment (SPS_{mhcz})

A key characteristic of J buyers in the model developed in Section 2 is that they trade with just one seller. Guided by this insight, we use the ratio of the number

[19] Including mode of transport in these bins mitigates the influence of spurious sources of variation like product quality that might differ across product varieties shipped using different methods.

[20] Quadruples with fewer than five shipments might also represent importers trying out a new product or other idiosyncrasies. In Appendix Section B, we provide some statistics comparing our sample against the broader sample of all arm's-length quadruples with at least two transactions. We need at least two transactions to be able to compute some of our variables, such as weeks between shipments (WBS_{mhcz}). As expected, the excluded quadruples with fewer than five transactions tend to be relatively small and trade more rarely.

[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. Moreover, we assume that A 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] Appendix C provides more details on how all variables are constructed. While below we also analyze quantity per shipment (QPS_{mhcz}) and unit value per shipment (UV_{mhcz}), they are not summarized here due to differences in the quantity units across products. We note that relationship lengths can be subject to both left and right censoring at the beginning and end of our 1992 to 2016 sample period.

Table 1: Attributes of $mhcz$ Quadruples

	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) quadruples during our 1992 to 2016 sample period. Observations are restricted to quadruples with at least five transactions. Observation counts are rounded to the nearest thousand per US Census Bureau disclosure guidelines.

of sellers to the number of shipments (SPS) within importer-product-country-mode ($mhcz$) quadruples,

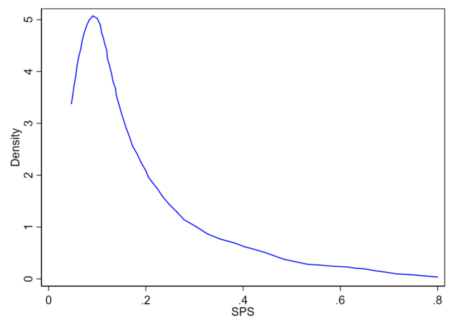
$$SPS_{mhcz} = \frac{Sellers_{mhcz}}{Shipments_{mhcz}}, \tag{6}$$

as an observable metric of J sourcing. This variable has an upper bound of one, i.e., a different supplier for every shipment, and approaches a lower bound of zero in the case of many transactions sourced from a single seller. Buyers that use fewer sellers relative to the number of shipments (i.e., those with lower values of SPS_{mhcz}) are more likely to be engaged in repeated transactions, and hence in J procurement. While A buyers might in theory also transact with few sellers if they repeatedly offer the lowest price, introducing noise into our measure, we find below that SPS_{mhcz} is indeed correlated with procurement patterns in a manner consistent with the model.

The distribution of SPS_{mhcz} across buyer quadruples with at least five transactions from 1992 to 2016 is displayed in the kernel density reported in Figure 3. As indicated in the figure, most buyer quadruples have a relatively small ratio of sellers to shipments. Observations in the right tail approach a value of 1, i.e., a different seller for each shipment. As reported in the final row of Table 1, the mean ratio of sellers to shipments across buyer quadruples is 0.33, with standard deviation 0.24.

The first two columns of Table 2 report the weighted average of SPS_{mhcz} for buyer quadruples trading with the noted countries, using the quadruples' total imports as weights. These means are reported for the two five-year time periods used in our

Figure 3: 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 all buyer quadruples with at least five transactions between 1992 and 2016. The figure was created according to Census Bureau guidelines and omits observations below the 5th percentile and above the 95th percentile.

regression analysis in Section Section 4. For the first time period, we find that the average SPS_{mhcz} is lowest for US imports from Mexico and Japan, consistent with the prevalence of J sourcing in the automobile industry—a key industry in US trade with these countries—including among large Japanese multinationals like Toyota (Boehm et al., 2020). Results in the second column reveal that, over time, average SPS_{mhcz} generally falls. The largest decreases exhibited, both in levels and percent growth, are for Mexico, China and Brazil. The relatively large drop for Mexico may be related to increasingly close supply-chain integration with US producers as a result of NAFTA. In Section 4, we examine whether the decline in SPS_{mhcz} for China is related to the US granting Permanent Normal Trade Relations (PNTR) in 2001.

In subsequent analysis, we will also consider an indicator variable for J importers that takes the value 1 when SPS_{mhcz} falls in the first quartile of its distribution computed within a given bin k in the first period, 1995-2000, J^k_{mhcz}. For our cross-country comparison, we compute the SPS distribution within product-mode bins, but across countries ($k = hz$). This choice implies that the share of J imports can vary between countries even though, worldwide, 25 percent of quadruples fall into the first quartile by construction. We define analogous dummies for the later time period, also with respect to the distribution of SPS_{mhcz} in the *first* time period, to capture changes with respect to the initial distribution. The final two columns of Table 2 report the share of imports from each country in each time period accounted for by buyer quadruples for which $J^{hz}_{mhcz} = 1$. While the 25^{th} percentile cutoff used in this procedure is arbitrary, it provides a rough indication of variation in J importing

Table 2: J Relationships by Country

	Mean SPS		$J_{mhcz}^{hz} = 1$ Share of Import Value	
	(1)	(2)	(3)	(4)
Country	1995-2000	2002-2007	1995-2000	2002-2007
Mexico	0.095	0.068	0.750	0.869
Japan	0.107	0.123	0.756	0.725
Taiwan	0.132	0.114	0.711	0.743
Canada	0.141	0.120	0.602	0.667
United Kingdom	0.146	0.225	0.717	0.519
South Korea	0.156	0.135	0.656	0.724
France	0.177	0.158	0.627	0.667
Rest of the World	*0.180*	*0.156*	*0.625*	*0.678*
Germany	0.184	0.163	0.582	0.606
China	0.185	0.147	0.582	0.693
Brazil	0.190	0.151	0.576	0.706

Source: LFTTD and authors' calculations. Columns (1) and (2) report the weighted average sellers per shipment (SPS_{mhcz}) across buyer quadruples with at least five transactions by country and period, where import values are used as weights. Columns (3) and (4) report the share of the value of US 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 column (1).

across source countries. Consistent with the raw SPS_{mhcz} measure, J import value shares increase over time, overall, and most strongly for Brazil, China, and Mexico.

Appendix B presents further breakdowns of how SPS_{mhcz} varies across groups of ten-digit HS codes and 6-digit NAICS industries of importing firms. We find that J sourcing is most prevalent for transportation equipment, machinery, and plastics, and that manufacturers are the most likely to use J sourcing, consistent with these firms obtaining relatively customized inputs for their production processes.

3.2 SPS_{mhcz} and Procurement Attributes

We now evaluate the link between SPS_{mhcz} and procurement patterns via an $mhcz$-level OLS regression,

$$\ln(\overline{Y}_{mhcz}) = \beta_1 \ln(SPS_{mhcz}) + \beta_2 \ln(QPW_{mhcz}) + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{hcz} + \epsilon_{mhcz}. \tag{7}$$

Guided by our theory, the dependent variable, $\overline{Y}_{mhcz}$, represents the key dimensions by which the A and the J systems differ: average quantity per shipment (QPS_{mhcz}), weeks between shipments (WBS_{mhcz}), and unit value (UV_{mhcz}) across all transactions within the $mhcz$ quadruple. In line with holding quantity fixed in Section 2, we condition on buyers' total order "flow" by controlling for the quantity imported by a buyer quadruple over its entire lifetime divided by its overall length, in weeks, QPW_{mhcz}.[23] We control for quadruples' first and last weeks of trade, beg_{mhcz} and end_{mhcz}, to capture effects of trading in a specific time period—such as a particular stage in the business cycle—and duration effects.[24] Our regression also includes product by country by mode of transportation fixed effects (λ_{hcz}), which capture time-invariant characteristics of trade along these dimensions such as distance, transit time, or level of transportation infrastructure. The sample period is 1992 to 2016, and standard errors are two-way clustered at the country and product level.[25]

Results for specification (7) are reported in Table 3. In the first three columns, we find that quadruples with higher SPS_{mhcz}, i.e., those that are more A, receive shipments for a given total order flow that are larger, less frequent, and lower in price, consistent with Proposition 2.3. Furthermore, the coefficient estimates for our quantity control, QPW_{mhcz}, are in line with Proposition 5.1, discussed below, where an increase in the total quantity procured leads to an increase in shipment size and reductions in the number of weeks between shipments and unit value. Coefficient estimates for SPS_{mhcz} indicate that increasing sellers per shipment by one standard deviation from its mean (from 0.33 to 0.58) is associated with a 23 log point rise in quantity per shipment, a 25 log point increase in weeks between shipments, and a 7 log point decline in price.[26]

In the final three columns of Table 3, we consider a related specification that

[23]This variable also controls for the possibility that overall order flow could lead to variation in average shipment sizes or unit values for reasons unrelated to the procurement system. 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.

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

[25]As before, we only use quadruples with at least five shipments over the entire sample period. In Appendix D, we show that results are qualitatively identical for a cutoff of 10 shipments. We describe the construction of all variables in detail in Appendix C.

[26]Our analysis computes 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 Appendix D, we re-run specification (7) where we compute SPS using the ratio of sellers to transactions within buyer-product-country triples (SPS_{mhc}) and buyer-product doubles (SPS_{mh}), and find similar results.

Table 3: SPS_{mhcz} and Procurement Attributes

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$
$\ln(SPS_{mhcz})$	0.418^{***}	0.452^{***}	-0.123^{***}			
	0.017	0.017	0.021			
$1\{SPS_{mhcz} = Q2\}$				0.328^{***}	0.350^{***}	-0.117^{***}
				0.014	0.015	0.014
$1\{SPS_{mhcz} = Q3\}$				0.552^{***}	0.591^{***}	-0.179^{***}
				0.024	0.024	0.023
$1\{SPS_{mhcz} = Q4\}$				0.792^{***}	0.856^{***}	-0.226^{***}
				0.034	0.035	0.038
$\ln(QPW_{mhcz})$	0.701^{***}	-0.308^{***}	-0.287^{***}	0.687^{***}	-0.323^{***}	-0.282^{***}
	0.014	0.014	0.020	0.013	0.014	0.019
Observations	2,966,000	2,966,000	2,966,000	2,966,000	2,966,000	2,966,000
Fixed effects	*hcz*	*hcz*	*hcz*	*hcz*	*hcz*	*hcz*
R-squared	0.947	0.674	0.845	0.945	0.661	0.845
Controls	beg, end	beg, end	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attributes of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}), sellers per shipment quartile dummies, and total quantity shipped per week (QPW_{mhcz}). (QPS_{mhcz}), (WBS_{mhcz}), and (UV_{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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

relaxes the restriction of a linear relationship between procurement attributes and sellers per shipment by replacing SPS_{mhcz} with a series of dummy variables indicating the quartile into which buyer quadruples fall. We compute these quartiles separately for each hcz bucket using the entire sample period. The first quartile, $1\{SPS_{mhcz} = Q1\}$, is the left-out category. This specification further justifies the use of SPS_{mhcz} as a metric of J sourcing, as coefficient estimates for SPS_{mhcz} rise or fall monotonically from quartile 1 to quartile 4 in a manner consistent with the quartiles representing increasingly A quadruples.[27]

[27] In Appendix Section D we show that the relationships displayed above are robust to analyzing procurement patterns separately by mode of transport, i.e., vessel versus air. In Supplemental Appendix L, on the authors' websites, we show that the results are similar when we examine procurement patterns within $mxhcz$ buyer-seller relationships, and that the results hold separately within each sector, such as manufacturing or retail. We also show that procurement patterns for differentiated products based on Rauch (1999) are more J compared to commodities.

Table 4: SPS_{mhcz} and Relationship Length

	(1)	(2)
Dep. var.	$\ln(length_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	-0.576^{***}	
	0.015	
$(SPS_{mhcz} = Q2)$		-0.383^{***}
		0.015
$(SPS_{mhcz} = Q3)$		-0.683^{***}
		0.027
$(SPS_{mhcz} = Q4)$		-1.139^{***}
		0.047
$\ln(QPW_{mhcz})$	-0.147^{***}	-0.130^{***}
	0.006	0.005
Observations	$2,966,000$	$2,966,000$
R-squared	0.431	0.413
Fixed effects	hcz	hcz
Controls	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing the average buyer-seller quadruple relationship length ($length_{mhcz}$) on quadruples' sellers per shipment (SPS_{mhcz}), sellers per shipment quartile dummies 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 quadruple, and exclude quadruples with less than 5 shipments. Standard errors, adjusted for clustering by country (c) and product (h) bin are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

3.3 SPS_{mhcz} and Other Characteristics

Relationship Length: Buyers under the J system rely on repeat purchases from the same seller, while buyers under the A system choose potentially different lowest-cost suppliers for each transaction. An implication of these choices is that J buyers have longer relationships with their suppliers. We investigate this prediction using the variable $length_{mhcz}$, which tracks the average length of the mx buyer-seller relationships associated with $mhcz$ buyer quadruples. This variable is constructed in two steps. First, for each $mxhcz$ quintuple, we compute the total number of weeks passed between the first and the last transaction of any product by any mode between the buyer m and seller x, i.e., their total relationship length. Second, for each $mhcz$ buyer quadruple, we take the average of these numbers of weeks across all $mxhcz$ quintuples within the quadruple. 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, add products over time.

We use the same specification outlined in equation (7) but using $length_{mhcz}$ as the dependent variable. The results, reported in Table 4, show that $mhcz$ buyer quadruples with lower ratios of SPS_{mhcz} tend to have longer relationships. In column (1), we find that a one standard deviation increase of sellers per shipment from its mean is associated with a 31 log point decrease in average relationship length. In column (2), the average relationship length for quadruples in the fourth quartile (most A) is about 114 log points lower than that in the first quartile (most J).[28]

Buyer Characteristics: In Appendix B, we show that the importer dimension is the most important for explaining variation in SPS_{mhcz}. We therefore next investigate how various firm-level attributes are related to import sourcing strategy, measured by SPS_{mhcz}. We aggregate the quadruples across products, countries, and modes to the importer-level and run the importer-level regression

$$\ln(\overline{Y}_m) = \beta_1 \ln(SPS_m) + \epsilon_{m\cdot}, \tag{8}$$

where $\overline{Y}_m$ is one of importer m's total sales, total payroll, average wage, or the firm's inventory-to-sales ratio, and SPS_m is an average of SPS_{mhcz} across all quadruples of the importer. We obtain sales, payroll, and wages at the firm-level from the Longitudinal Business Database (LBD), where the average wage is constructed as the firm's total payroll divided by the number of employees. We obtain beginning-of-year inventories for manufacturing firms from the Annual Survey of Manufactures (ASM) and the Census of Manufactures (CMF). We use for each firm attribute the earliest non-missing observation available for the firm.[29] Table 5 shows the regression results.

We find that firms that on average rely on more A procurement practices tend to be smaller, pay lower wages, and hold higher inventories. An increase in average sellers per shipment by one standard deviation from its mean is associated with a 16 log point decline in sales, 19 log point decline in payroll, and a 3 log point decline in the average wage.[30] A one standard deviation increase in SPS_{mhcz} from its mean raises the inventory-to-sales ratio by 0.8 log points, consistent with A procurement leading to larger inventories.[31]

[28]In Appendix section D and Supplemental Appendix L, we show that all our robustness checks also go through for the length variable. In Supplemental Appendix L, we also consider an alternative definition of relationship length where we treat each quintuple as a separate relationship, rather than using the overall importer-supplier pair, and show that our results still hold.

[29]Results are robust to using an average across all active years (see Appendix D).

[30]As we will show in Section 6 below, these findings are qualitatively consistent with our model, where we find that larger importers are more likely to use the J system.

[31]Note that we only observe the overall inventories of the firm, across all products and including

Table 5: SPS_m and Firm Characteristics

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(sales_m)$	$\ln(pay_m)$	$\ln(wage_m)$	$(inv/sales)_m$
$\ln(SPS_m)$	-0.291^{***}	-0.350^{***}	-0.056^{***}	0.015^{***}
	0.005	0.006	0.002	0.001
Observations	184,000	184,000	184,000	48,500
R-squared	0.015	0.018	0.003	0.006

Source: LFTTD and authors' calculations. Table reports the results of regressing importer characteristics in the year of the importer's first transaction on sellers per shipment (SPS_{mhcz}) averaged across all quadruples involving the importer. All regressions exclude quadruples with less than five shipments. ($sales_m$), (pay_m), ($wage_m$), and (($inv/sales)_m$) are total sales, total payroll, average wage (i.e., payroll divided by number of employees), and total inventory at the beginning of the year divided by total sales, respectively. Robust standard errors are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Finally, consistent with a firm-wide sourcing strategy, we find that importers' procurement system is correlated across products. Using all importers with at least two products in a given country-mode bin, we randomly draw two of these products for each importer. We then use the J indicator J^{hcz}_{mhcz}, computed using the distribution of SPS_{mhcz} within hcz bins for the entire sample period, and regress $J^{hcz}_{mhcz,1}$ of the first product on $J^{hcz}_{mhcz,2}$ of the second product. We re-run this regression 1000 times, where we re-draw the two products on every run. Our estimated average coefficient on $J^{hcz}_{mhcz,2}$ is 0.234 (bootstrap s.e. = 0.001) with a constant of 0.214 (s.e. = 0.001), indicating that the probability of the second product being J approximately doubles when the first one is.

4 PNTR and the Choice of Procurement System

A key insight from the model presented in Section 2 is that trade policy can alter buyers' choice of procurement system by affecting the probability of trade wars. In this section, we examine the prediction that a decrease in the probability of a trade war can induce buyers to shift from A to J procurement using a plausibly exogenous change in US trade policy, the US granting of Permanent Normal Trade Relations (PNTR) to China in 2001. We assess these shifts across both continuing and new $mxhcz$ quintuples, and in terms of importers' sellers per shipment (SPS).

domestic purchases. Our results suggest that variation in international procurement is associated with tangible differences in firms' overall inventories.

As described in Pierce and Schott (2016), prior to PNTR, US imports from China were subject to the risk of punitive tariff increases absent annual action from the President and Congress. Pierce and Schott (2016) and Alessandria et al. (2024) document the trade-dampening effects of this uncertainty on US importers prior to PNTR, and Handley and Limão (2017) provide a theoretical basis for these effects that operates via suppressed entry by Chinese exporters. We measure exposure to PNTR via the "NTR Gap" from Pierce and Schott (2016), which measures the amount that tariffs could have increased prior to PNTR and varies by product.[32]

PNTR and continuing mxhcz quintuple attributes: Our first approach to testing whether PNTR influences procurement is to examine its impact on the procurement attributes examined in Section 3: quantity per shipment, weeks between shipments and unit value. These attributes are observed at the buyer-seller $mxhcz$ quintuple level. We therefore analyze procurement attributes among *continuing* quintuples, which trade in both the pre- and the post-PNTR period, in this subsection, and for new quintuples in the next subsection.

Our OLS triple difference-in-differences (DID) identification strategy examines the relationship between PNTR and the procurement attributes before versus after the change in policy (first difference), for imports from China versus other source countries (second difference), for products with higher versus lower NTR gaps (third difference),

$$\begin{aligned} \ln(Y_{mxhczt}) = {} & \beta_1 1\{t = Post\} * 1\{c = China\} * NTR\,Gap_h \\ & + \beta_2 ln(QPW_{mxhczt}) + \beta_3 \chi_{mxhczt} + \lambda_{mxhcz} + \lambda_t + \epsilon_{mxhczt}. \end{aligned} \tag{9}$$

The last difference captures the fact that products with larger NTR gaps experience a greater decline in the relationship termination probability, which is a function of the change in China's NTR status (identical for all products) and the increase in tariff rates that could have occurred before PNTR, which varies by product. We expect the largest shift toward J procurement after PNTR to occur in US imports of high-NTR-gap products from China.

The variable Y_{mxhczt} on the left-hand side of specification (9) represents one of the

[32] See Supplemental Appendix M, on the authors' websites, for details on the NTR gap variable. While the probability that tariff increases would occur was identical across products, the probability of such an increase severing importer-supplier relationships varies with the NTR Gap.

Table 6: Baseline Within $mxhcz$ Quintuple PNTR DID Regression

	(1)	(2)	(3)
Dep. var.	$\ln(QPS_{mxhczt})$	$\ln(WBS_{mxhczt})$	$\ln(UV_{mxhczt})$
$Post_t * China_c * NTRGap_h$	-0.197***	-0.168***	0.092***
	0.009	0.009	0.023
$ln(QPW_{mxhczt})$	0.368***	-0.632***	-0.124***
	0.009	0.008	0.013
Observations	439,000	439,000	439,000
R-squared	0.982	0.894	0.985
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 2002 to 2007. QPS_{mxhczt}, WBS_{mxhczt}, and UV_{mxhczt} are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period t. All regressions include $mxhcz$ and period t fixed effects, control for the beginning and end week of the quadruple as well as all variables needed to identify the DID term of interest. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

three procurement attributes: quantity per shipment (QPS_{mxhczt}), weeks between shipments (WBS_{mxhczt}), and unit value (UV_{mxhczt}).[33] The first term on the right-hand side is the triple difference-in-differences (DID) term of interest, an interaction of an indicator for the post period, $1\{t = Post\}$, a dummy for imports from China, $1\{c = China\}$, and the $NTR\ Gap_h$. The variable χ_{mxhczt} 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_{mxhczt}) as well as buyer-seller quintuple (λ_{mxhcz}) and period (λ_t) fixed effects. Our two five-year periods (t), 1995 to 2000 and 2002 to 2007, are chosen to straddle the change in policy in 2001 and end before the Great Recession.[34] Standard errors are two-way clustered at the country and product level.

Conducted at the $mxhcz$ level, equation (9) is restricted to continuing buyer-seller relationships via the $mxhcz$ quintuple fixed effect. We restrict the sample to quintuples that transact at least twice both before PNTR and after the policy change so that weeks between shipments (WBS_{mxhczt}) can be computed.

Results, reported in Table 6, indicate that higher exposure to PNTR is associated

[33] Appendix C provides more details on how the variables in this section are constructed.

[34] In Appendix Section E, we demonstrate that all results in this section are robust to using a different post-PNTR period, 2004 to 2009.

with changes in shipping attributes that are consistent with a move 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 $NTR\ Gap$ (0.23) induces a relative decline in quantity per shipment and weeks between shipments of 4.5 log points and 3.9 log points respectively. Moreover, results in column 3 reveal that a one standard deviation increase in exposure to PNTR is associated with a relative increase in unit value of 2.1 log points. In each case, the findings in Table 6 are consistent with the predictions of Propositions 2.1 and 2.3, indicating a switch from A to J procurement, as opposed to an adjustment within the J system.[35]

PNTR and new mxhcz quintuple attributes: We next compare the procurement attributes of *new* buyer-seller $mxhcz$ quintuples formed in the post-PNTR period to relationships that were new in the pre-PNTR period. For both periods, we define new quintuples as those involving buyer-seller mx pairs that had not yet appeared before the beginning of the period, i.e., from 1992 to 1994 for the first period and from 1992 to 2001 for the second period.

As in the previous section, the regression is performed at the $mxhcz$ level and standard errors are two-way clustered at the country and product level. Instead of $mxhcz$ quintuple fixed effects, however, we include separate buyer quadruple ($mhcz$), exporter (x), and period (t) fixed effects, thereby focusing on buyers and sellers that exist in both time periods (with at least one trading partner), but who form new relationships across the time periods.[36]

Results, reported in Table 7, are consistent with relatively greater entry of J relationships after PNTR: buyer-seller quintuples trading goods with greater exposure to the change in policy formed after it was implemented exhibit relatively smaller and more frequent shipments, at relatively higher prices, than quintuples formed before PNTR. Point estimates indicate that a one standard deviation increase in exposure is associated with a 2.7 log point and 2.2 log point decline in shipment size and frequency, respectively, and a 2.1 log point rise in price.

[35] Consistent with Proposition 5.1, the coefficient estimates for $ln(QPW_{mxhczt})$ indicate that an increase in the procurement quantity increases the size of shipments, raises shipping frequency, and reduces unit values. We show in Appendix E that our conclusions are qualitatively unchanged, though the coefficient on WBS_{mxhcz} is not statistically significant, when we do not include QPW_{mxhczt} as a covariate. In Supplemental Appendix N, we analyze the effect of PNTR on the three procurement attributes at the $mhcz$ quadruple level and find similar results.

[36] As noted in Supplemental Appendix N, results are robust to including both continuing and new $mxhcz$ buyer-seller quintuples simultaneously in one regression.

Table 7: New $mxhcz$ Quintuple PNTR DID Regression

	(1)	(2)	(3)
Dep. var.	$\ln(QPS_{mxhczt})$	$\ln(WBS_{mxhczt})$	$\ln(UV_{mxhczt})$
$Post_t * China_c * NTRGap_h$	-0.116***	-0.097***	0.090**
	0.023	0.023	0.038
$ln(QPW_{mxhczt})$	0.409***	-0.594***	-0.129***
	0.013	0.012	0.018
Observations	3,184,000	3,184,000	3,184,000
R-squared	0.966	0.842	0.972
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 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_{mxhczt}), (WBS_{mxhczt}), and (UV_{mxhczt}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period t. All regressions include $mhcz$, x and period t fixed effects, control for the beginning and end week of the quadruple as well as all variables needed to identify the DID term of interest. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

PNTR and Sellers per Shipment (SPS): The previous two exercises demonstrate that higher exposure to PNTR is associated with relatively more J procurement attributes after the policy change. We next focus on the impact of PNTR on buyers' sellers per shipment, the metric for identifying J relationships introduced in Section 3. We consider both the continuous measure SPS_{mhcz} as well as the indicator for whether this ratio falls into the first quartile of the pre-PNTR period distribution within product by country by mode bins, $J^{hcz}_{mhcz} = 1$.

Our triple DID specification is similar to equation (9), but takes place at the buyer $mhcz$ quadruple level,

$$\begin{aligned} ln(Y_{mhczt}) =& \beta_1 1\{t = Post\} * 1\{c = China\} * NTRGap_h + \beta_2 ln(QPW_{mhczt}) + \\ & \beta_3 \chi_{mhczt} + \lambda_{mhcz} + \lambda_t + \epsilon_{mhczt}. \end{aligned} \quad (10)$$

The triple DID term of interest is the same as above, an interaction of post-period and China-import dummies with the NTR gap, and the variable χ_{mhczt} 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_{mhczt}) as well as buyer quadruple (λ_{mhcz}) and period (λ_t) fixed effects. Once again, standard errors are two-way clustered at the country and

Table 8: Within-Importer PNTR Regression, Buyer Characteristics

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(SPS_{mhczt})$	$1\{J^{hcz}_{mhczt}=1\}$	$\ln(SPS_{hczt})$	$1\{J^{hcz}_{hczt}=1\}$
$Post_t * China_c * NTRGap_h$	-0.006	0.041*	-0.021**	0.034*
	0.031	0.022	0.009	0.019
$ln(QPW_{mhczt})$	-0.171***	0.124***	-0.062***	0.032***
	0.006	0.005	0.002	0.003
Observations	738,000	291,000	368,000	28,500
R-squared	0.772	0.675	0.695	0.547
Fixed effects	$mhcz, t$	$mhcz, t$	hcz, t	hcz, t
Controls	Yes	Yes	Yes	Yes

Source: LFTTD and authors' calculations. First two columns report 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. Second two columns are analogous but at the hcz level of aggregation. Pre- and post-PNTR periods are 1995 to 2000 and 2002 to 2007. All regressions include period t fixed effects, control for the beginning and end week of the quadruple as well as all variables needed to identify the DID term of interest. Columns 2 and 4 exclude quadruples with less than five shipments in both periods. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

product level. Conducted at the $mhcz$ level, equation (10) is restricted to continuing importers—i.e. those active before and after granting of PNTR—via the $mhcz$ buyer quadruple fixed effect.

While our model requires repeated interactions between buyers and sellers, it does not mandate relationships be long-established. Moreover, existing research finds substantial relative growth in US-importer-Chinese-exporter relationships after PNTR (Pierce and Schott, 2016). As a result, we also estimate equation (10) at the more aggregated hcz level, which broadens the analysis to include entering importers. For this regression, SPS_{hcz} is defined as the average SPS_{mhcz} within hcz cells. The J indicator J^{hcz}_{hcz} is defined analogously.

Results in Table 8 indicate that PNTR induced a shift towards more J importing, with this effect driven by the entry of new importers. As shown in the first two columns, we find no relationship between the policy change and SPS_{mhczt}, though continuing buyer quadruples more exposed to the change in policy exhibit relative increases in the probability of being in the first SPS quartile ($J^{hcz}_{mhczt} = 1$) after PNTR. When we re-estimate equation (10) at the hcz level—which accounts for the role of new importers that enter in the post-PNTR period—we find that higher exposure to PNTR is associated not only with a higher probability of $J^{hcz}_{mhczt} = 1$, but also with a precisely estimated reduction in SPS_{hczt}. The estimates in columns 3 and 4 indicate that a one standard deviation increase in exposure to PNTR is associated with a 0.5 percent relative decline in SPS and a 1.3 percent increase in falling within the first quartile of the SPS distribution.[37]

Overall, the results in this section provide support for the model's prediction that a lower likelihood of a trade war can bring buyers to switch from A to J procurement in terms of the procurement attributes with importers' partners, the formation of new J relationships, and the number of seller partners.

5 Multi-Country Setup with Endogenous Demand

In this section, we embed the partial equilibrium model introduced in Section 2 within the multi-country, multi-product general equilibrium model of Eaton and Kortum (2002). We use the model to assess the potential welfare implications of shutting down J procurement in Section 7. Such analysis is of particular relevance given recent efforts to reverse globalization, such as "Brexit" and the Trump trade wars, that have increased trade policy uncertainty worldwide.

Our main point of departure from Eaton and Kortum (2002) is the introduction of homogeneous *buyer* firms in each country, which purchase manufacturing goods from *sellers* and distribute these goods to consumers. Buyer and seller firms are subject to the procurement problem described in Section 2.

5.1 Environment

Households: Our modeling is standard. There are N countries, indexed by n and i. Each country is populated by L_n consumers, who purchase a continuous flow of a manufactured composite good and a homogeneous "outside" good 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 good. The composite good is a CES aggregate of a continuum of differentiated products indexed by $\omega \in$

[37] To analyze the influence of initial buyer experimentation during the years immediately after PNTR on our results, we also consider, in Appendix E, 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.

$[0, 1]$,

$$Q_n = \left(\int_0^1 q_n(\omega)^{(\sigma-1)/\sigma} d\omega \right)^{\sigma/(\sigma-1)}, \tag{11}$$

where $\sigma > 0$ is the elasticity of substitution and $q_n(\omega)$ is quantity. This aggregator implies the standard price index $P_n = \left(\int p_n(\omega)^{1-\sigma} d\omega \right)^{1/(1-\sigma)}$. We assume that each consumer supplies one unit of labor.

Homogeneous good: The homogeneous good in country n is produced by a representative firm according to $Z_n = a_n L_n^O$, where a_n is productivity and L_n^O is the aggregate labor used in the production of the good. Labor is paid the wage rate w_n. The homogeneous good is directly sold to households and can be costlessly traded across countries. We set its price as the numeraire and normalize it to one. Labor is perfectly mobile between manufacturing and the homogeneous good sector, as in Eaton and Kortum (2002), Handley and Limão (2017), and Antràs et al. (2017).

Manufacturing sellers: Manufactured good ω can be produced by homogeneous seller firms in country n with the linear production function $q = \frac{\Upsilon}{\theta} l$ introduced in Section 2. Sellers are perfect competitors, taking prices as given. Their productivity $\Upsilon_n(\omega)$ is specific to each origin country-product combination. Sellers in country n incur fixed logistic and transport costs f_n in units of seller country labor for each destination supplied. We assume that a country's firms are owned by their households.

Manufacturing buyers: We add a continuum of homogeneous buyer firms in each country into the standard framework. Buyer firms purchase manufactured goods from sellers domestically or abroad, and offer them to the households in their country at prices $p_n(\omega)$. The transactions between buyers and sellers take place as described in Section 2: given household demand $q_n(\omega)$, buyers in country n choose the lowest-cost sourcing country i for product ω, the procurement system, and the optimal order size. Buyers using the A system need to use an additional $m_n(\omega)$ labor units to inspect the quality of the good. Buyers choosing the J system pay an incentive premium to ensure quality.[38] As discussed in Section 2, due to the fixed procurement cost each buyer optimally places each order with a single seller.

[38]The incentive payments imply that sellers obtain positive profits under the J system. These profits are not competed away since sellers offering a lower price would violate the incentive constraint.

5.2 Partial Equilibrium with Endogenous q

In this section, we describe how $q_n(\omega)$ is determined in equilibrium. We assume that the buyer has already chosen the source country and procurement system and discuss how these are chosen in the next section. As we are focusing on a single market in this section, we omit country and product subscripts.

Our first step to determine the equilibrium, Proposition 5.1, establishes that batch size and shipping frequency increase with quantity ordered, q:

Proposition 5.1. *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.5. □

Intuitively, for a given fixed shipping frequency, buyers must increase the batch size x_s in both systems to meet an increase in q. But by the first-order condition (5), buyers 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. As a result, larger quantities lead to both greater shipment sizes, x_s, but also greater order frequencies. Unit values decrease since fixed per-shipment costs are spread over greater per-shipment quantities. Additionally, in the J 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 3 and 4. As indicated in Tables 3, 6, and 7, we find that shipment size is positively related to the quantity shipped per week (QPW), and that both weeks between shipments and unit values are negatively related to QPW.

We next show that buyers' average cost curves are downward sloping:

Lemma 5.2. *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. Appendix A.6, shows that average cost curves are downward sloping. Supplemental Appendix J.2 shows that they are convex and converge to a finite limit. □

In our model, sellers face the standard constant marginal costs and perfect competition, but the fixed logistic and transport costs generate a natural monopoly for buyers in the downstream market. Downward sloping average cost curves are a key departure of our model from trade models based on Eaton and Kortum (2002), which generally assume constant marginal cost. We therefore need to choose an appropriate market structure. Our assumption is that buyers compete in a "contestable" market for consumers, a natural extension of Bertrand competition when firms' costs exhibit economies of scale (Baumol et al., 1982, Tirole, 1988). In a contestable market, there exist several homogeneous competitors whose entry is costless. Due to downward sloping average costs, in equilibrium a single buyer serves the entire consumer market for each product.

Lemma 5.2 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. If the buyer were to price above average costs, entrants would contest the positive profits and take over the market. If the buyer were to price below average costs, she would realize negative profits. Since for any $q < q^*$ consumers are willing to pay prices greater than average costs, potential entry forces an incumbent offering q to lower its prices and to increase quantity to the level q^* where supply equals demand.

Under appropriate assumptions on the demand system, the market equilibrium is a corollary of Lemma 5.2.[39]

Corollary 5.2.1. *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.*

[39] In principle our CES demand system may intersect the downward sloping average cost curve multiple times. For equilibrium to exist in that case, the demand curve must cut the average cost curve from above at the intersection that determines the greatest equilibrium quantity, q^*_{high}. Intuitively, if the demand curve were to cut from below, it would be above the average cost curve for all $q^*_{high} < q \to \infty$, implying that consumers are willing to buy an infinite quantity of the good when the buyer sets price equal to average costs.

5.3 General Equilibrium with Endogenous q

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

Cost minimization: Buyer firms in country n minimize average costs $AC_n(q_n(\omega))$ of purchasing $q_n(\omega)$ by choosing the lowest-cost system and country:

$$AC_n(q_n(\omega))^* = \min\left\{\min\left\{AC_{ni,A}(x^*_{ni,A}(\omega), q_n(\omega)), AC_{ni,J}(x^*_{ni,J}(\omega), q_n(\omega))\right\};\ i = 1, ..., N\right\}, \tag{12}$$

where $AC_{ni,s}(x^*_{ni,s}(\omega), q_n(\omega))$ are average costs of purchasing $q_n(\omega)$ under system s from country i, and $x^*_{ni,s}(\omega)$ is the optimal batch size determined by the first-order condition (5). Since average costs are downward sloping in q and the market is contestable, in equilibrium there is only one buyer firm serving each market. The contestable market price is $p_n(\omega) = AC_n(q_n(\omega))^*$.

Utility maximization: Consumption of each manufactured good is chosen to maximize (11) subject to the budget constraint

$$\int_0^1 p_n(\omega)q_n(\omega)\, d\omega \leq \alpha\left(w_nL_n + \sum_i\sum_s\int \pi^s_{in,s}(\omega)I_{in,s}(\omega)d\omega\right). \tag{13}$$

The right-hand side of the equation is the share of country n's total income W_n spent on manufacturing goods. Since labor is perfectly mobile between sectors, the wage rate is pinned down by the productivity of the homogeneous good sector as $w_n = a_n$. The second term on the right-hand side, which is new relative to the standard framework, represents the incentive premia collected from shipments to countries i under $s = J$. Here, $\pi^s_{in,s}(\omega)$ is the continuous flow of profits to sellers in country n from sales to country i of product ω under system s, and $I_{in,s}(\omega)$ is an indicator that is equal to one if seller n uses system s to country i. Profits are zero if shipments are under the A system. Consumption of the homogeneous good satisfies $Z_n = (1-\alpha)W_n$.

Market clearing: Equilibrium requires market clearing for each manufactured good

ω and for the homogeneous good, and labor market clearing in each country. We provide these market clearing conditions in Appendix F.[40]

6 Quantitative Analysis

In this section, we estimate the model quantitatively before using it in Section 7 to analyze the effects of changes in trade policy on trade flows and welfare. This analysis highlights the impact of firms' choice of procurement system on the welfare gains from trade, as well as the relevance of the model for the current international trading environment. We parametrize the model using a combination of external calibration and within-model moment matching.

Due to the non-linearity of the buyer's problem, our model does not admit an analytical solution. We therefore use an iterative algorithm. First, given parameter values, we compute the average cost curves for each market and system. Next, we guess each country's price index and income to compute the demand curve in each market and find the last intersection where demand intersects the lowest average cost curve from above. Given this equilibrium in each market, we compute a new price index and income in each country, construct a new demand curve, and iterate to convergence. Appendix G provides further details.

6.1 Parametrization and Calibrated Parameters

We set each time period to one quarter. We set $N = 3$ countries and interpret these countries to be the United States, China, and the Rest of the World (RoW).[41] As in Eaton and Kortum (2002), productivity $\Upsilon_n(\omega)$ is drawn from a Fréchet distribution $F_n(\Upsilon) = e^{-\Lambda_n \Upsilon^{-\zeta}}$, where the country-specific parameter Λ_n scales the mean of the distribution and ζ scales the variation. The productivity draws are independent across products within each country.

We assume inspection costs for domestic procurement to be zero, implying that all domestic sourcing takes place under the A system.[42] For imports, we assume that the

[40]In the quantitative simulations, we verify that a positive amount of labor is allocated to both manufacturing and the production of the homogeneous good in each country in equilibrium.

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

[42]This is a normalization. Since we do not have domestic transactions data, we cannot estimate

distribution of inspection costs is Pareto and given by $G_n(m) = 1 - (\underline{m}/m)^{\gamma_n}$, where $\underline{m}$ is the lower bound of inspection costs, and γ_n is a parameter to be estimated.[43] We set $\underline{m} = 0.001$ to reflect the fact that inspection is essentially costless for some goods, e.g., commodities. Heterogeneous inspection costs generate dispersion in the relative costs of A and J procurement, and hence in the system used, across goods coming from the same country. The shape of the inspection cost distribution is directly tied to the welfare effects of policy: if some products have more extreme inspection costs, then a high probability of trade war that forces firms to use the A system for these products can lead to large welfare losses.

We calibrate a number of parameters outside of the model, and summarize their values in Table 9. We provide more information on the calibration in Appendix H, and discuss here only the rate of exogenous relationship break-ups, ρ_{ni}. In the model, this variable reflects any exogenous shock that ends relationships. We assume that this break-up rate is symmetric between country pairs, $\rho_{ni} = \rho_{in}$, and set it for the US by fitting the exponential decay parameter that best matches the empirically observed fraction of plausibly J buyer-seller ($mxhcz$) quintuples that survive for $2, ..., 100$ quarters in the US trade data. Since trade wars between the United States and the RoW are unlikely in steady state, we interpret the estimated decay parameter for relationships between US and RoW firms, equal to 0.087, as normal churn due to firm exits, product obsolescence, and so on.[44] We therefore set $\rho_{US,RoW} = 0$. For relationships between US and Chinese suppliers, we estimate a decay parameter of 0.114. We interpret this higher likelihood of break-ups as arising due to the additional uncertainty of trading with China, and thus set the relationship break-up rate between the US and China equal to the difference in the decay parameters, leading to $\rho_{US,CN} = 0.0264$. For trade between China and RoW, we set $\rho_{CN,RoW} = 0$ as well.[45]

the share of J and A trade for within-country transactions. Equivalently, we could also refer to "domestic" sourcing as a third type of procurement system that does not face an incentive problem and hence corresponds to the first-best outcome.

[43]We perform an alternative estimation below where we assume that inspection costs are distributed according to a Fréchet distribution instead of Pareto. We found that the model fit is better under Pareto, and therefore choose it as our baseline. We also assume that a given destination country has the same distribution of inspection costs for all origins to reduce the degrees of freedom in the estimation. We show below that the model fits the data quite well despite this restriction.

[44]While narrow trade disputes between the United States and RoW—such as safeguards and antidumping duties—occur often, the WTO's formal dispute settlement system was an effective deterrent to full-fledged trade war between the US and RoW during our sample period.

[45]While trade tensions were also present between RoW and China, a variety of bilateral disagreements between the US and China meant that the risk of RoW-China trade war was substantially

Table 9: Calibrated Parameters

Parameter	Value	Source
Interest rate (r)	0.01	Caliendo et al. (2019)
Elasticity of substitution (σ)	3.85	Antràs et al. (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)
Homogeneous good sector productivity (a_n)		
- US	1	Normalization
- China	0.12	World Bank, authors' calculations
- RoW	0.47	World Bank, authors' calculations
Labor Force (L_n)		
- US	1	Normalization
- China	5	World Bank, authors' calculations
- RoW	2.5	World Bank, authors' calculations
Rate of exogenous break-ups, US -China ($\rho_{US,CN}$)	0.0264	Census Bureau (LFTTD)
Rate of exogenous break-ups, US -RoW ($\rho_{US,RoW}$)	0	Assumption

Notes: Table presents the exogenously fixed parameters. Column (1) displays the parameter value, and column (2) shows its source.

Appendix H provides more details.

6.2 Targeted Moments and Estimation

We estimate the remaining productivity scales T_n, the country-specific fixed costs f_n, and the inspection cost distribution parameters γ_n via simulated method of moments using the LFTTD and aggregate data. The column labeled "Moment in Data" in Table 10 summarizes the values of the targeted moments in the data. We next describe the empirical moments targeted and the underlying identification assumptions. Appendix H provides more details.

We normalize $T_{US} = 1$, and estimate the other two productivity parameters using the share of imports from China and from the rest of the world in US domestic manufacturing sales in 2016 (rows 1 and 2 of Table 10). A lower value of T_n increases country n's productivity, which raises that country's share in US domestic sales.

We estimate the remaining four parameters using the observed shipping patterns in the trade data. A corollary of Proposition 2.2 is that, given a total quantity ordered q, higher fixed costs lead to shipments that are less frequent under both systems. We can therefore estimate the fixed shipment costs f_{CN} and f_{RoW} by running a modified

lower.

Table 10: Estimated Parameters and Targeted Moments

	(1)	(2)	(3)	(4)	(5)
	Parameter	Estimated Value	Moment that Primarily Identifies the Parameter	Moment in Data	Moment in Model
(1)	Productivity China (T_{CN})	15.482	Share of Chinese imports in domestic sales	0.074	0.066
(2)	Productivity RoW (T_{RoW})	2.745	Share of RoW imports in domestic sales	0.270	0.276
(3)	Fixed costs, China (f_{CN})	0.310	$\exp(\hat{\beta}_0 + \hat{\beta}_1 + \hat{\beta}_3\overline{beg} + \hat{\beta}_4\overline{end})$ from (14) for CN	91.00	91.10
(4)	Fixed costs, RoW (f_{RoW})	0.061	$\exp(\hat{\beta}_0 + \hat{\beta}_1 + \hat{\beta}_3\overline{beg} + \hat{\beta}_4\overline{end})$ from (14) for RoW	60.90	62.66
(5)	Dispersion of inspection	0.290	$\hat{\beta}_1$ from (14) for China	0.871	0.814
(6)	costs, China (γ_{CN})		Sd of $\hat{\epsilon}$ from (14) for China	0.227	0.180
(7)	Dispersion of inspection	0.101	$\hat{\beta}_1$ from (14) for RoW	0.822	0.818
(8)	costs, RoW (γ_{RoW})		Sd of $\hat{\epsilon}$ from (14) for RoW	0.219	0.207
(9)	Total objective $T(\cdot)$				0.062

Source: LFTTD and authors' calculations. Column (1) lists the parameters estimated for the model. Column (2) contains the estimated parameter values. Column (3) reports the moment targeted to identify the parameter. Column (4) presents the value of the moment in the data, and Column (5) presents the value of the moment computed in our simulated model. The last row presents the value of the function $T(\cdot)$ from (15).

classification regression (7) with average weeks between shipments ($\overline{WBS}_{mhcz}$) as dependent variable, separately for China and for the rest of the world,

$$\begin{aligned} \ln(\overline{WBS}_{mhcz}) = \beta_0 + \beta_1 1\{\overline{WBS}_{mhcz} = Q4\} + \beta_2 \ln(QPW_{mhcz}) \\ + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{hcz} + \epsilon_{mhcz}. \end{aligned} \tag{14}$$

We control for the total quantity per week, QPW, to be consistent with the theory, and for time variation and fixed effects by product by country by mode to remove potentially confounding variation that is unrelated to fixed costs. To isolate sourcing that is most likely under the A and the J system, our regression sample includes only quadruples that fall into the first or the fourth quartile of the SPS_{mhcz} distribution (hence, are most likely J and A, respectively), and includes a dummy, $1\{\overline{WBS}_{mhcz} = Q4\}$, indicating whether $\overline{WBS}_{mhcz}$ falls into the fourth quartile. We set f_n by targeting the predicted average shipping frequency in the fourth quartile, $\exp(\hat{\beta}_0+\hat{\beta}_1+\hat{\beta}_3\overline{beg}+\hat{\beta}_4\overline{end})$, where $\overline{beg}$ and $\overline{end}$ are the simple averages of beg_{mhcz} and end_{mhcz} in the data (rows 3 and 4).[46] Since we do not have information on the procurement choice of foreign importers sourcing from the US, we assume $f_{US} = f_{RoW}$.

[46]Since quantity units are very heterogeneous across goods in the data, we target the shipping frequency at $\ln(QPW_{mhcz}) = 0$. We target the average shipping frequency within the fourth quartile, hence likely A procurement, to remove variation in shipping patterns that is due to different mixes of A versus J sourcing.

Regression (14) is also informative about the dispersion of the inspection cost parameters γ_{CN} and γ_{RoW}, which are crucial for the share of J sourcing estimated by the model. Starting from $\gamma_n \to \infty$, at which point all inspection costs are zero and all sourcing is under the A system, lowering γ_n increases the number of high inspection cost draws and therefore raises the share of J sourcing. We target two sets of moments that we obtain from regression (14). First, we target the difference in shipping frequencies between the first and the fourth quartile of the WBS_{mhcz} distribution, given by $\hat{\beta}_1$ in specification (14) (rows 5 and 7). A greater dispersion of inspection costs (a smaller γ_n) increases the difference in average shipping times between those quadruples that are more likely A and those that are more likely J. Second, we target the dispersion in shipping times across more A $mhcz$ quadruples. When γ_n is low, the inspection cost draws are more dispersed, leading to a higher variance of the shipping frequencies within the A system. We construct this moment by taking the residuals from (14) for all observations that fall into the fourth quartile of the WBS distribution, and compute the standard deviation of these residuals. We generate the moments in exactly the same way in the model.[47] We prefer this approach to the alternative of simply setting the shares of A and J sourcing exogenously. Rows 6 and 8 show the estimated moments. Similar to the fixed costs, we assume that $\gamma_{US} = \gamma_{RoW}$.

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

$$\phi^* = \arg\min_{\phi \in \mathbb{F}} \sum_x T(\mathcal{M}_x(\phi), \hat{\mathcal{M}}_x) \tag{15}$$

where $T(\cdot)$ is the percentage difference between the model, $\mathcal{M}_x(\phi)$, and data, $\hat{\mathcal{M}}_x$, moments. Appendix I.1 provides more details on the estimation algorithm and outcomes. We present the estimated values of the parameters in the column labeled "Estimated Value" in Table 10, and the "Moment in Model" column shows the values of the simulated moments with these parameters.

The model provides a good fit along several dimensions. First, the model-generated shares of Chinese and RoW imports in US manufacturing consumption are 6.6% and 27.6%, respectively, compared to 7.4% and 27.0% in the data. Second, the model generates shipping frequencies consistent with the data: the time between shipments

[47] We do not include beg_{mhcz}, end_{mhcz}, and the fixed effects λ_{hcz} in regression (14) run in the model since there are no changes over time and the random parameter values are drawn from the same stationary distribution for all products.

is about 91 weeks for China and 63 weeks for the rest of the world, conditional on $\ln(QPW_{mhcz}) = 0$.[48] Finally, the model generates substantial variation in shipping frequencies across goods, similar to the data. Our results slightly underestimate the dispersion of inspection costs for China (rows 5-6). Increasing the dispersion of inspection costs further would raise the average time between shipments beyond its empirical target (row 3), but would tend to increase the share of J sourcing.

The fixed costs of production in terms of labor are about five times larger for China than for the rest of the world (rows 3 and 4). Since wages in China are four times lower than in RoW, the fixed costs in terms of the numeraire good are only slightly higher (about 20 percent). These higher fixed costs are an implication of the lower shipping frequency from China compared to the rest of the world. Since the estimation target includes the intercept β_0, which is estimated using the observed trade flows in our sample period, the higher fixed cost reflects any trade barriers between countries in our sample, such as distance (Hummels and Schaur, 2013).

6.3 Model Results

The first column of Table 11 summarizes 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 J system. Our estimates imply that 10 percent of imports from China are under the J system, while 52 percent of imports from the rest of the world take place under J procurement. The higher share for the rest of the world reflects the higher trade war probability with China, which discourages trade under the J system, as well as the higher fixed costs for China, which makes the frequent shipments under the J system more expensive. The structurally estimated J shares are somewhat smaller than the empirical estimates we obtained using shipments in the first quartile of the SPS_{mhcz} distribution in Table 2 for China, but they are in the ballpark for the rest of the world.

Rows 6 to 7 of Table 11 show that the average product imported by the US is subject to an inspection cost of 0.4 percent and a fixed cost of 4.5 percent of the import value, respectively. These figures provide a validity check of the model, since

[48] The empirically observed number of weeks between shipments is much lower since shipping frequency increases with quantity shipped. In the first quartile of the WBS distribution from China the average number of weeks between shipments is 9 weeks, in the fourth quartile it is 39 weeks.

Table 11: Comparison of Base-Model and Counterfactual Equilibria

		(1)	(2)	(3)	(4)
		Baseline Equilibrium	**Equilibrium Without Japanese Sourcing**	**Autarky**	**Removal of PNTR**
(1)	Share of consumption from China (%)	6.6%	7.1%	.	6.6%
(2)	- of which, J	9.5%	.	.	7.1%
(3)	Share of consumption from ROW (%)	27.6%	19.6%	.	27.6%
(4)	- of which, J	52.1%	.	.	52.1%
(5)	Share of consumption from US (%)	65.8%	73.3%	100.0%	65.8%
(6)	Avg. inspection costs	0.4%	1.3%	.	0.4%
(7)	Avg. fixed costs (imports)	4.5%	3.3%	.	4.5%
(8)	Manufacturing price index	1.000	1.029	1.122	1.000
(9)	Utility	1.000	0.982	0.940	0.9998

Table shows various statistics of the equilibrium under the assumption of a Pareto distribution for inspection costs. The first column presents the statistics for the baseline equilibrium, using the parameters that minimize the objective function. The second column shows the same statistics for a counterfactual economy in which the formation of J relationships is not possible due to $\rho \rightarrow \infty$. The third column shows an autarky economy in which trade is not possible. The fourth column shows a counterfactual economy in which we reduce the arrival rate of trade wars from China to zero. Rows 1-5 show the share of US manufacturing sales, $P_{US}Q_{US}$, that is from China, from the rest of the world, and from the US , respectively, and the share of these manufacturing sales that is sourced under the J system. Row 6 presents the average inspection costs as a share of the import value, computed over all imports, including under the J system. Row 7 shows the average fixed costs as a share of the import value. Row 8 shows the manufacturing price index, P_{US}, normalized to one in the baseline. Row 9 shows total utility, $W_{US} = Q_{US}^{\alpha} Z_{US}^{1-\alpha}$, normalized to one in the baseline.

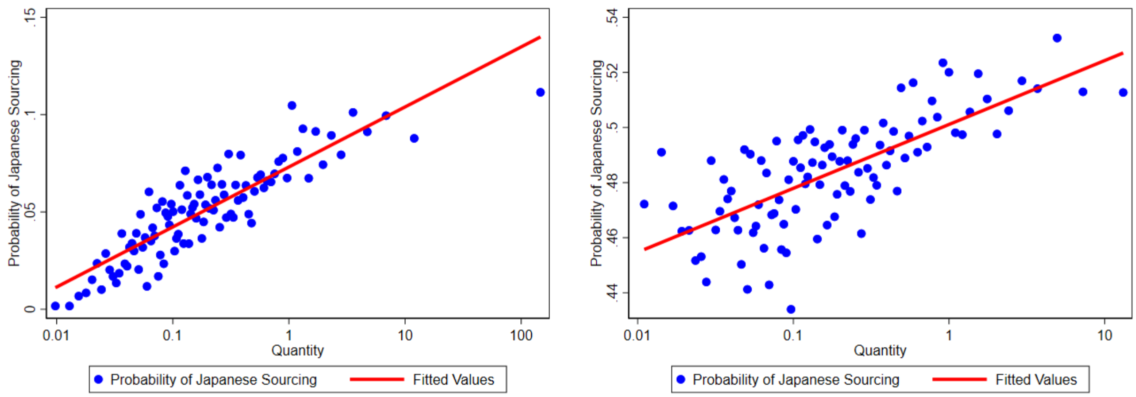
they are in line with estimates by Kropf and Sauré (2014), who estimate that Swiss exporters face total fixed shipment costs of 0.8 percent to 5.4 percent of the value imported. The final two rows present the price index in manufacturing in the United States, P_{US}, and the utility $Q_{US}^{\alpha} Z_{US}^{1-\alpha}$, normalized to 1 for ease of interpretation.

As a further check of the model, we verify that larger importers are more likely to use the J system, as found in Table 5 above. We plot in Figure 4 the average share of J importers against the average quantity imported for each percentile of the quantity distribution of imports, for China and RoW.[49] The figure shows that larger importers are more likely to use the J system, as in the data. Intuitively, a higher seller productivity raises imports under both systems by reducing variable costs. Under the J system a higher seller productivity additionally lowers the incentive premium, which makes J sourcing relatively more attractive for high-productivity imports.

[49] We drop outliers below the 1st and above the 99th percentile of the distribution.

Figure 4: Quantity Imported vs Share of J Importers

(a) China (b) Rest of the World



Notes: The figure shows for each percentile of the distribution of US imports the average quantity imported against the average share of importers using the J system. The left panel presents the results for imports from China, the right panel is for imports from the rest of the world.

7 Counterfactuals

In this section, we use the calibrated model to estimate US welfare under two counterfactuals. First, to determine the importance of J sourcing, we compare elimination of J sourcing *vis à vis* autarky. This counterfactual then provides context for estimating the welfare gain associated with PNTR. Each of these analyses is highly relevant for considering the effects of recent increases in uncertainty in the global trading system.

No Japanese Sourcing: We shut down J importing by setting $\rho_{US,n} = \infty$ for trade between the US and both of its trading partners. As shown in the second column of Table 11, US imports rise slightly from 6.6 to 7.1 percent for China (row 1), while imports fall significantly from 27.6 to 19.6 percent from RoW (row 3).[50] Intuitively, buyers' procurement choice in our model is shaped by three factors: (i) seller productivity, (ii) country-product inspection costs, and (iii) the country-pair probability of trade peace. Products in which the domestic country has a high productivity are sourced domestically. Products not sourced domestically are imported under the A system if inspection costs are low, and under the J system if they are high and a trade war is unlikely. A greater likelihood of trade war raises the incentive premia under the J system, rendering A and domestic sourcing more attractive. As the arrival rate

[50]This exercise entails a relatively larger increase in trade costs for RoW than for China since in the baseline the trade war arrival rate with the rest of the world is zero while it is 0.0264 for China. As a result, there is a relative shift of trade towards China in this counterfactual.

of trade wars goes to infinity, no goods are imported under the J system.

Table 11 shows that average inspection costs jump from 0.4 to 1.3 percent (row 6) because the imports which switch from J to A are precisely those with relatively high inspection costs, for which A sourcing was previously not optimal. Higher import prices drive down the average fixed cost as a share of import value (row 7). The manufacturing price index P_{US} rises 2.5 percent (row 8) due to higher sourcing costs. Overall, welfare falls by 1.6 percent. This drop in welfare is about one third as severe as moving the US to autarky, as illustrated in the third column of Table 11.

In Appendix I.2, we check the robustness of these findings by re-estimating the model with a Fréchet rather than Pareto distribution for inspection costs. This estimation matches our targeted moments slightly less well than the baseline, but generates significantly higher import shares under the J system. We find that welfare costs of removing J relationships are significantly larger, at 3.5 percent, indicating that the losses rise substantially with the share of J relationships.[51]

Removal of PNTR: Our second counterfactual, summarized in the fourth column of Table 11, analyzes a hypothetical scenario in which PNTR is removed. Alessandria et al. (2022) estimate an annual probability that NTR is revoked between 8 and 15 percent in the 1990s, and Handley and Limão (2022) find a probability of around 13 percent. We take this latter estimate since it falls within the range of Alessandria et al. (2022). We increase $\rho_{US,CN}$ from the baseline relationship break-up rate of 0.0264, which implies that relationships break with about 10 percent probability over four quarters, by 13 percentage points so that it implies a break-up rate of 23 percent over four quarters, leading to $\rho_{US,CN} = 0.0654$.[52] We then re-simulate the model.

As indicated in the last column of the table, we find that the share of J imports from China decreases by 2.5 percentage points as a result of the higher possibility of relationship break-ups, and the share of imports from China falls by 0.1 percentage point. The overall price and welfare effects are very minor, leading to a welfare loss of 0.02 percent.

Our results differ significantly from Handley and Limão (2017), who find larger effects of PNTR on consumer income, for several reasons. First, in our model changing

[51]In Supplemental Appendix O, we illustrate the effect of changing $\rho_{US,n}$ for intermediate values between zero and infinity on the share of US consumption, welfare, and consumer income.

[52]While our baseline $\rho_{US,CN}$ is computed using the entire sample period, the post-PNTR break-up rate is very similar, $\rho^{post}_{US,CN} = 0.022$. Hence, when we add 13 percentage points to this number the results are very similar.

the probability of a trade war only affects products imported under the J system, which account for less than one tenth of consumer expenditures on Chinese goods. Importers do not bear the full costs of the trade war but can switch to the A system, which mutes the increase in costs. Our exercise highlights that the welfare costs of a trade war could be significantly higher if a trade war affects countries with a high share of J relationships, such as RoW, or additionally impedes contract enforcement under the A system. Second, in the Handley-Limão model a reduction of uncertainty leads to the entry of foreign exporters, thus expanding the set of available varieties and driving down the price index, as in the Melitz (2003) model. In contrast, while in our framework a change in trade policy uncertainty may change the identity of the supplier of a good, the set of available varieties is fixed as in Eaton and Kortum (2002). We view our channel as complementary to the mechanisms described by Handley and Limão (2017).[53]

Discussion: The counterfactuals considered in this section, though stylized, highlight the potential importance of relational contracting in the welfare gains from trade. They also demonstrate that the firm and country losses associated with greater trade policy uncertainty depend on the choice of procurement system, and therefore the costs associated with switching systems. Firms (and countries) that disproportionately import hard-to-inspect goods that make greater use of the J system when trade wars are unlikely will experience the largest welfare losses when uncertainty rises, as switching to the A system will be most costly. We note that the welfare losses implied by our framework likely capture only a fraction of the true losses associated with greater trade policy uncertainty because our framework considers trade only in final goods, and just 34 percent of US consumption is imported. In reality, many of the 66 percent of domestically produced consumption goods contain imported inputs and would therefore also be susceptible cost increases as trade wars break out. We leave modeling this channel to future research.

[53] Handley and Limão (2017) also allow exporters to pay a fixed cost to reduce their marginal cost, and the set of firms that choose to pay this cost rises when uncertainty is lower, which further increases the gains from low uncertainty they find.

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' solution to a quality control problem depends upon exporters' beliefs about the possibility of a trade war between the firms' countries. When the probability of trade war is high, buyers choose "American"-style procurement, characterized by large, infrequent orders and costly inspections. When the probability of trade war is low, buyers can induce sellers to provide high quality by paying a premium over a long-term relationship. We show that changes in trade policy can induce a switch between procurement systems.

We examine the model's key implications using transaction-level US import data, and show that importer-exporter relationships differ along the dimensions – such as shipment size, shipment frequency and unit value – emphasized in the model. Using a triple difference-in-differences specification, we also show that PNTR is associated with a movement toward more Japanese-style procurement among US importers and Chinese exporters along the dimensions highlighted by the model. Quantitative simulations reveal that an increase in the probability of trade war that is sufficient to eliminate "Japanese"-style procurement reduces US welfare about one third as much as placing the US in autarky.

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.[54] 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.

[54]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.

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Online Appendix

A Analytical Results

A.1 Effect of Quality and Trade Wars on Average Costs

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

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

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

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

A.2 Proof of Proposition 2.1

For $\bar{\theta} - \underline{\theta} > 0$ and $\rho > 0$, when $m_A = 0$ average costs under the J system must be higher than under the A system by the discussion above Proposition 2.1 and in Appendix A.1. Since average costs under the A system grow without bound as $m_A \to \infty$, there must be an m^* such that average costs under the systems are equalized.

A.3 Proof of Proposition 2.2

Japanese System: We apply the implicit function theorem to the FOC (5):

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

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

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

where we denote by SOC_J the second-order condition, which is greater than zero as shown in Supplemental Appendix J.1.

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:

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

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.4 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 as shown in Supplemental Appendix J.1.

American System

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

Japanese System

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

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

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

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 $y \equiv r x_J^*/q$, where $0 < y < 1$. We find for the right-hand side $\lim_{y\to 0}\frac{y e^{-y}}{1 - e^{-y}} = \lim_{y\to 0} 1 - y = 1$. Next, note that $\frac{d}{dy}\frac{y e^{-y}}{1 - e^{-y}} = \frac{e^{-y}\left[(1-y) - e^{-y}\right]}{\left[1 - e^{-y}\right]^2} < 0$. It follows that the right-hand side of (A.1) 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: $v_A(x_A)/x_A < v_J(x_J)/x_J$

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

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

A.5 Proof of Proposition 6.1

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

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

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.6 Proof of Lemma 6.2: Average cost curves are downward sloping

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

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

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

$$AC'(q) = \frac{-\frac{f+m}{q^2}}{1 - exp(-\frac{rx}{q})}. \tag{A.3}$$

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

$$AC'(q) = \frac{-\frac{f}{q^2}}{1 - exp(-\frac{rx}{q})}. \tag{A.4}$$

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 and converge to a finite limit. See Supplemental Appendix J.2, available on the authors' websites.

B Data Refinement and Summary Statistics

We use version c201601 of the LFTTD data, which we refine as follows. First, we drop all transactions that are warehouse entries. Second, we remove all transactions that do not include a valid importer 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, we exclude from our analysis all related-party transactions.[55] We choose a conservative approach and exclude all relationships in which the two parties ever report being related, as well as 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 HS10 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.[56] 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. 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, 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.

Our baseline sample restricts our cleaned data to importer (m) by HS10 product (h) by country (c) by mode of transportation (z) $mhcz$ quadruples with at least five transactions. Table A.1 provides some details for our sample period 1992-2016. We compare this sample to an alternative arm's-length sample that does not restrict to buyer quadruples with at least five transactions in Supplemental Appendix K.

Table A.2 provides information on the average number of sellers per shipment (SPS_{mhcz}) by ten-digit HS code, analogous to Table 2 in the main text. For columns (3) and (4), we define J dummies J^k_{mhcz} that take a value of one if SPS_{mhcz} falls in the first quartile of its distribution within country-mode bins in the first time period ($k = cz$) to retain variation across products. We find that J sourcing is most prevalent for transportation equipment, machinery, plastics, and optical products. We show a similar table by the main 6-digit NAICS industry of the importer in Supplemental Appendix K, and show that manufacturers are the most likely to use J sourcing.

Most of the variation in SPS_{mhcz} is driven by importers. We run a series of regressions of SPS_{mhcz} separately on importer, product, country, importer industry, and mode of transportation fixed effects, and examine the R-squared from these regressions to study how much of the variation is explained.[57] We find that importer,

[55] 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.

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

[57] For industry, we use 6-digit NAICS fixed effects. We define the importer's main industry in

product, industry, country, and mode fixed effects individually explain 35%, 12%, 10%, 8%, and 7% of the variation in SPS_{mhcz}, respectively. The large heterogeneity in SPS_{mhcz} across importers is consistent with different firms choosing different procurement strategies.

Table A.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
Unique 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. Observations are restricted to quadruples with at least five transactions. Import values are in billions of real 2009 dollars. Vessel imports refer to imports arriving over water. The final four rows of the table 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. Observation counts are rounded to the nearest thousand per U.S. Census Bureau disclosure guidelines.

Table A.2: "Japanese" Relationships by HS Category

	Mean SPS		$J^{cz}_{mhcz}=1$ Share of Import Value	
	(1)	(2)	(3)	(4)
Product code (HS chapter)	1995-2000	2002-2007	1995-2000	2002-2007
Transportation (86-89)	0.107	0.081	0.783	0.880
Machinery (84-85)	0.130	0.133	0.754	0.763
Plastics (39-40)	0.130	0.096	0.727	0.820
Optical products (90-92)	0.137	0.127	0.726	0.768
Footwear (64-67)	0.142	0.117	0.750	0.827
Other products (93-99)	*0.151*	*0.124*	*0.697*	*0.808*
Metals (72-83)	0.154	0.128	0.600	0.737
Food (16-24)	0.155	0.120	0.601	0.747
Chemicals (28-38)	0.156	0.121	0.600	0.736
Stones & Jewelry (68-71)	0.159	0.141	0.658	0.674
Animal products & vegetables (01-15)	0.166	0.132	0.511	0.608
Minerals (25-27)	0.182	0.203	0.570	0.500
Leather and wood products (41-49)	0.188	0.153	0.556	0.688
Textiles (50-63)	0.224	0.177	0.463	0.604

Source: LFTTD and authors' calculations. The first two columns report the weighted average sellers per shipment (SPS_{mhcz}) across buyer quadruples with at least five transactions by HS category and period, where import values are used as weights. Numbers in parentheses refer to the Harmonized System chapter of the product. The second two columns report the share of the value of US imports accounted for by quadruples with SPS_{mhcz} in the first quartile of the distribution of SPS_{mhcz} within country-mode in the first period. Rows of the table are sorted by column (1).

each year as the one with the largest share of employment, and then take the modal main industry across the years in which the quadruple is active.

C Construction of the Variables

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" (i) refers to a week in which the quadruple imports. Table A.3 provides a summary of how we construct the variables in Section 3. Table A.4 describes the variables used in Section 4.

Table A.3: Classification Regressions

	Formula	Description
Quantity per Shipment (QPS_{mhcz})	$\frac{\sum_i Quantity_{mhczi}}{Ntrans_{mhcz}}$	$Quantity_{mhczi}$ is the quantity imported by quadruple $mhcz$ at transaction i and $Ntrans_{mhcz}$ is the total number of transactions by the quadruple in 1992-2016.
Weeks between Shipments (WBS_{mhcz})	$\frac{end_{mhcz}-beg_{mhcz}}{Ntrans_{mhcz}-1}$	end_{mhcz} is the number of the week of the last transaction of the quadruple and beg_{mhcz} is the number of the week of the first transaction of the quadruple in t (see definition below). The denominator represents the number of time periods between subsequent transactions of the quadruple, which is one less than the number of transactions. If $Ntrans_{mhczi}=1$, the average time gap cannot be computed.
Unit Value (UV_{mhcz})	$\frac{1}{Ntrans_{mhcz}}\sum_i \frac{Value_{mhczi}}{Quantity_{mhczi}}$	$Value_{mhczi}$ is the value imported by quadruple $mhcz$ at transaction i, $Quantity_{mhczi}$ is the corresponding quantity.
Quantity per Week (QPW_{mhcz})	$\frac{\sum_i Quantity_{mhczi}}{end_{mhcz}-beg_{mhcz}}$	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 five transactions in our baseline, the beginning and end week are never the same and therefore the expression is finite.
First week (beg_{mhcz}) Last week (end_{mhcz})	$min\{Week_{mhczi}\}$ $max\{Week_{mhczi}\}$	$Week_{mhczi}$ 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.
Avg. relationship length ($length_{mhcz}$)	$\frac{\sum_x length_{mx}}{Sellers_{mhcz}}$	$length_{mx}=max\{Week_{mxi}\}-min\{Week_{mxi}\}$. $Week_{mxi}$ is the week number of a transaction i of the buyer-seller pair mx in any good or mode of transportation, relative to the first week of 1960. $Sellers_{mhcz}$ is the number of exporters (x) with which the quadruple ($mhcz$) has an $mxhcz$ quintuple relationship.

Table A.4: PNTR Regressions

	Formula	Description
Quantity per Shipment (QPS_{mxhczt})	$\frac{\sum_i Quantity_{mxhczti}}{Ntrans_{mxhczt}}$	$Quantity_{mxhczti}$ is the quantity imported by quintuple $mxhcz$ in period t (either 1995-2000 or 2002-2007) at transaction i and $Ntrans_{mxhczt}$ is the total number of transactions by the quintuple in period t.
Weeks between Shipments (WBS_{mxhczt})	$\frac{end_{mxhczt}-beg_{mxhczt}}{Ntrans_{mxhczt}-1}$	end_{mxhczt} is the number of the week of the last transaction of the quintuple in period t (either 1995-2000 or 2002-2007) and beg_{mxhczt} is the number of the week of the first transaction of the quintuple. The week number is relative to the first week of 1960. Thus, for example the first week of 2016 has week number 2912. The denominator represents the number of time periods between subsequent transactions of the quintuple, which is one less than the number of transactions. If $Ntrans_{mxhczt} = 1$, the average time gap cannot be computed. The PNTR regressions therefore require for each quintuple at least two transactions in each period t.
Unit Value (UV_{mxhczt})	$\frac{1}{Ntrans_{mxhczt}} \sum_i \frac{Value_{mxhczti}}{Quantity_{mxhczti}}$	$Value_{mxhczti}$ is the value imported by quintuple $mxhczt$ at transaction i in period t, and $Quantity_{mxhczti}$ is the corresponding quantity.
Quantity per Week (QPW_{mxhczt})	$\frac{\sum_i Quantity_{mxhczti}}{end_{mxhczt}-beg_{mxhczt}}$	In contrast to QPS_{mxhczt}, 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_{mxhczt}, we require for each quintuple at least two transactions in each period t so that this variable can be computed.

D Additional A vs J Classification Regressions

Thicker Relationships: Our baseline regressions in Section 3.2 are restricted to $mhcz$ quadruples with at least five transactions over our sample period. One concern might be that for quadruples that trade only relatively few times, our variable suppliers per shipment (SPS_{mhcz}) is mismeasured because we did not observe a sufficient number of transactions. In Table A.5, we show that our results are robust to restricting the regression to quadruples with at least 10 transactions.

More Aggregated Suppliers per Shipment: Another concern with our measure of SPS 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 or even mh level. In Tables A.6 and A.7 we show that our results are robust to defining SPS at these higher levels of aggregation (i.e., SPS_{mhc} or SPS_{mh}), where we keep all other variables at the $mhcz$ level of the baseline.

Different Modes of Transportation: We next investigate whether the results hold separately for vessel vs air shipments. Results in Table A.8 indicate similar results for both forms of transport.

Average Firm Attributes: In regression (8), we use the firm-level attribute in the year of the firm's first import transaction. In Table A.9 we instead compute for each buyer quadruple an average of the firm attribute across all years in which the quadruple is active, and then average across quadruples. The two specifications could generate different results if the firm's attributes change significantly over time. The results are similar to the baseline.

Table A.5: A vs J Classification Regression With At Least 10 Transactions

	(1)	(2)	(3)	(4)
Dep. var.	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$	$\log(length_{mhcz})$
$\log(SPS_{mhcz})$	0.359***	0.370***	−0.064***	−0.504***
	0.015	0.016	0.020	0.013
$\log(QPW_{mhcz})$	0.700***	−0.306***	−0.273***	−0.134***
	0.014	0.014	0.019	0.005
Observations	1,645,000	1,645,000	1,645,000	1,645,000
R-squared	0.950	0.659	0.855	0.488
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 (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). QPS_{mhcz}, WBS_{mhcz}, P_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, 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 quadruple, and exclude quadruples with less than 10 shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.6: A vs J Classification Regression With SPS at mhc Level

	(1)	(2)	(3)	(4)
Dep. var.	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$	$\log(length_{mhcz})$
$\log(SPS_{mhc})$	0.346***	0.376***	−0.083***	−0.578***
	0.014	0.015	0.018	0.013
$\log(QPW_{mhcz})$	0.687***	−0.322***	−0.279***	−0.147***
	0.015	0.015	0.020	0.005
Observations	2,966,000	2,966,000	2,966,000	2,966,000
R-squared	0.944	0.654	0.844	0.442
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}, P_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.7: A vs J Classification Regression With SPS at mh Level

	(1)	(2)	(3)	(4)
Dep. var.	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$	$\log(Length_{mhcz})$
$\log(SPS_{mh})$	0.285***	0.311***	−0.063***	−0.483***
	0.019	0.020	0.021	0.009
$\log(QPW_{mhcz})$	0.668***	−0.343***	−0.274***	−0.115***
	0.014	0.014	0.020	0.006
Observations	2,966,000	2,966,000	2,966,000	2,966,000
R-squared	0.940	0.631	0.844	0.379
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}, P_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.8: A vs J Classification Regression Across Mode of Transport

	(1)	(2)	(3)	(4)
Dep. var.	$\log(QPS_{mhcz})$	$\log(WBS_{mhcz})$	$\log(UV_{mhcz})$	$\log(length_{mhcz})$
	Vessel			
$\log(SPS_{mhcz})$	0.419***	0.451***	-0.172***	-0.570***
	0.015	0.015	0.013	0.018
$\log(QPW_{mhcz})$	0.661***	-0.347***	-0.263***	-0.177***
	0.011	0.011	0.018	0.008
Observations	1,506,000	1,506,000	1,506,000	1,506,000
R-squared	0.924	0.686	0.829	0.434
	Air			
$\log(SPS_{mhcz})$	0.410***	0.443***	-0.058**	-0.609***
	0.022	0.022	0.025	0.018
$\log(QPW_{mhcz})$	0.737***	-0.272***	-0.300***	-0.106***
	0.015	0.015	0.023	0.005
Observations	1,029,000	1,029,000	1,029,000	1,029,000
R-squared	0.933	0.635	0.764	0.416

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}, 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.9: SPS_m and Firm Characteristics

	(1)	(2)	(3)	(4)
Dep. var.	$\log(sales_m)$	$\log(pay_m)$	$\log(wage_m)$	$(inv/sales)_m$
$\log(SPS_m)$	-0.255^{***}	-0.313^{***}	-0.066^{***}	0.016^{***}
	0.005	0.006	0.002	0.001
Observations	184,000	184,000	184,000	48,500
R-squared	0.012	0.014	0.004	0.007

Source: LFTTD and authors' calculations. Table reports the results of regressing importer characteristics averaged across all years in which the importer is active on sellers per shipment (SPS_{mhcz}) averaged across all quadruples involving the importer. All regressions exclude quadruples with less than five shipments. ($sales_m$), (pay_m), ($wage_m$), and (($inv/sales)_m$) are total sales, total payroll, average wage (i.e., payroll divided by number of employees), and total inventory at the beginning of the year divided by total sales, respectively. Robust standard errors are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

E Additional DID Regressions

Alternate Time Periods: We show that our baseline DID results also hold if we use a different post-PNTR period from 2004 to 2009. Table A.10 presents the results from the continuing relationship PNTR regression (9), and Table A.11 shows the results for the regression with only new relationships. All results retain their expected sign and remain significant. Table A.12 presents the results from the within-importer regression, equation (10), both at the *mhcz* level and at the *hcz* level. On average, we find that the results from the main text become stronger for this later post-period, possibly because the shift of systems takes time.

No Quantity Control: 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. We therefore run the baseline PNTR regression (9) without quantity control, QPW_{mxhczt}. Results in Table A.13 show that we still find a decline in the quantity per shipment and an increase in the unit value. The effect on weeks between shipments is qualitatively consistent with the theory, though not significant at conventional levels.

Table A.10: Within $mxhcz$ Quintuple PNTR DID Regression: 2004-2009 vs 1995-2000

	(1)	(2)	(3)
Dep. var.	$\ln(QPS_{mxhczt})$	$\ln(WBS_{mxhczt})$	$\ln(UV_{mxhczt})$
$Post_t * China_c * NTRGap_h$	-0.199***	-0.163***	0.149***
	0.017	0.021	0.031
$ln(QPW_{mxhczt})$	0.403***	-0.606***	-0.133***
	0.009	0.008	0.014
Observations	221,000	221,000	221,000
R-squared	0.980	0.883	0.982
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_{mxhczt}), (WBS_{mxhczt}), and (UV_{mxhczt}) are average quantity per shipment, average weeks between shipments, and average unit value (i.e. value divided by quantity) in period t. All regressions include $mxhcz$ and period t fixed effects, control for the beginning and end week of the quintuple as well as all variables needed to identify the DID term of interest. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.11: New $mxhcz$ Quintuple PNTR DID Regression: 2004-2009 vs 1995-2000

	(1)	(2)	(3)
Dep. var.	$\ln(QPS_{mxhczt})$	$\ln(WBS_{mxhczt})$	$\ln(UV_{mxhczt})$
$Post_t * China_c * NTRGap_h$	-0.087**	-0.067*	0.075*
	0.036	0.035	0.045
$\ln(QPW_{mxhczt})$	0.414***	-0.590***	-0.127***
	0.012	0.011	0.017
Observations	3,158,000	3,158,000	3,158,000
R-squared	0.968	0.845	0.973
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_{mxhczt}), (WBS_{mxhczt}), and (UV_{mxhczt}) are average quantity per shipment, average weeks between shipments, and average unit value (i.e. value divided by quantity) in period t. All regressions include $mxhcz$ and period t fixed effects, control for the beginning and end week of the quintuple as well as all variables needed to identify the DID term of interest. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.12: Within-Importer PNTR Regression: 2004-2009 vs 1995-2000

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(SPS_{mhczt})$	$1\{J^{hcz}_{mhczt} = 1\}$	$\ln(SPS_{hczt})$	$1\{J^{hcz}_{hczt} = 1\}$
$Post_t * China_c * NTRGap_h$	-0.076**	0.076**	-0.027**	0.042
	0.037	0.029	0.011	0.027
$ln(QPW_{mhczt})$	-0.186***	0.125***	-0.059***	0.031***
	0.005	0.005	0.002	0.004
Observations	556,000	225,000	355,000	28,000
R-squared	0.757	0.660	0.687	0.550
Fixed effects	$mhcz, t$	$mhcz, t$	hcz, t	hcz, t
Controls	Yes	Yes	Yes	Yes

Source: LFTTD and authors' calculations. First two columns report 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. Second two columns are analogous but at the hcz level of aggregation. Pre- and post-PNTR periods are 1995 to 2000 and 2004 to 2009. All regressions include period t fixed effects, and control for the beginning and end week of the quadruple as well as all variables needed to identify the DID term of interest. Regressions in columns two and four are restricted to quadruples with at least five transactions in both periods. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table A.13: Baseline Within $mxhcz$ Quintuple PNTR DID Regression Without Quantity: 2002-2007 vs 1995-2000

	(1)	(2)	(3)
Dep. var.	$\ln(QPS_{mxhczt})$	$\ln(WBS_{mxhczt})$	$\ln(UV_{mxhczt})$
$Post_t * China_c * NTRGap_h$	-0.2753***	-0.0339	0.1186***
	0.0076	0.0318	0.0191
Observations	439,000	439,000	439,000
R-squared	0.97	0.69	0.98
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 2002 to 2007. (QPS_{mxhczt}), (WBS_{mxhczt}), and (UV_{mxhczt}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period t. All regressions include $mxhcz$ and period t fixed effects, control for the beginning and end week of the quadruple as well as all variables needed to identify the DID term of interest. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

F Market Clearing Conditions

Goods market clearing implies that production equals consumption for each ω:

$$\sum_n \sum_i \sum_s I_{ni,s}(\omega) x^*_{ni,s}(\omega) = \sum_n \sum_i \sum_s I_{ni,s}(\omega) \int_0^{x^*_{ni,s}(\omega)/q_n(\omega)} q_n(\omega) dt \quad \forall \omega, \quad \text{(A.5)}$$

where $I_{ni,s}(\omega)$ is an indicator function that is equal to one if the buyer in country n procures product ω from country i under system s, and zero otherwise.

The market for the homogeneous good clears as well,

$$\sum_n Z_n = \sum_n a_n L^O_n. \quad \text{(A.6)}$$

Finally, labor market clearing in each country requires that

$$L_n = \sum_i \sum_s \int_0^1 I_{in,s}(\omega) \frac{\bar{\theta}}{\Upsilon_n(\omega)} q_i(\omega) d\omega + f_n \sum_i \sum_s \int_0^1 I_{in,s}(\omega) \frac{q_i(\omega)}{x^*_{in,s}(\omega)} d\omega$$
$$+ \sum_i \int_0^1 I_{ni,A}(\omega) m(\omega) \frac{q_n(\omega)}{x^*_{ni,s}(\omega)} d\omega + L^O_n \quad \forall n \in N, \quad \text{(A.7)}$$

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 the homogeneous "outside" good 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 discretize the product space to $\Omega = 5,000$ products, and follow the steps in Table A.14. Our algorithm first computes the average cost curves and shipment sizes on a grid of inspection costs, productivities, trade war arrival rates, and quantities. We then guess a price index and total income for each country, trace out the demand curves, find the intersection of supply and demand, and iterate to convergence. We compute the average cost curves outside of the iteration algorithm since the numerical solution of the buyer's problem is quite time consuming. While in principle it would be possible to solve the buyer's problem within each iteration for each $ni\omega$ tuple, using linear interpolation on a grid during the iteration process is much faster.

Table A.14: Equilibrium Solution Algorithm

Step	Description
1	Initiate the model by drawing an inspection cost $m(\omega)$ for each product ω and country n from $G_n(m)$ and by drawing a productivity $\Upsilon_n(\omega)$ from $F_n(\Upsilon)$. Also set the trade war arrival rates ρ_{ni} for each country pair.
2	Define a four-dimensional grid with $(K_1 \times K_2 \times K_3 \times Q)$ grid points, where $K_1 = 70$, $K_2 = 60$, $K_3 = 60$, and $Q = 70$. 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 (4), i.e. $AC_A(\mathbf{k}) = \min_x \left(\frac{r}{q_k}\right) \frac{k_1 + k_2 x}{\left[1 - e^{-rx/q_k}\right]}$ and $AC_J(\mathbf{k}) = \min_x \left(\frac{r}{q_k}\right) \frac{k_1 + e^{(r+k_3)x/q_k} k_2 x}{\left[1 - e^{-rx/q_k}\right]}$. We denote by $x_A(\mathbf{k})$ and $x_J(\mathbf{k})$ the cost-minimizing shipment sizes under each system at grid point $\mathbf{k}$.
3	Map the draw $(m(\omega), \Upsilon_i(\omega), \rho_{ni})$ of each origin country (i)-destination country (n)-product (ω) triplet to an estimated average cost for each q_k using linear interpolation on the grid of average costs computed in Step 2, where under the A 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 J 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, $x^*_{ni,s}$, from linear interpolation on the grid of shipment sizes computed in Step 2.
4	Determine the cost minimizing system and origin country at each quantity q_k for each destination-product market $n\omega$, using equation (12). This traces out the average cost curve $AC_{n\omega}(q_k)$ of each market.
5	Begin iteration $t = 0$. Guess an initial manufacturing price index in each destination country, $P_n(t)$, and an initial total income, $W_n(t)$.
5.a	Compute each destination-product market $n\omega$'s demand curve, using utility maximization, by computing for each q_k the price $p_n(\omega; q_k, t) = \left(\frac{\alpha W_n(t)}{q_k}\right)^{\frac{1}{\sigma}} P_n(t)^{\frac{\sigma-1}{\sigma}}$.
5.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. Using the equilibrium prices in each market, compute a new price index, $P_n(t+1)$.
5.c	Determine the labor used for production, fixed costs, and inspection costs. Use the labor market clearing condition (A.7) to determine labor used for the homogeneous good sector L^O_n. Verify that this labor is non-negative.
5.d	Compute the total income in each country, $W_n(t+1)$, which is equal to labor income $w_n L_n$ plus profits under the "Japanese" system, see equation (13). Return to Step 5.a with $\{P_n(t+1), W_n(t+1)\}$ and iterate to convergence.

H Parameters and Empirical Moments

Table A.15 provides more detail on how we set the calibrated parameters in Table 9. Table A.16 contains more detail on how we construct the moments for the estimation.

Table A.15: Calibrated Parameters

<table>
<tr><th>Parameter</th><th>Description</th></tr>
<tr><td>Interest rate (r)</td><td>As in Caliendo et al. (2019)</td></tr>
<tr><td>Elasticity of substitution (σ)</td><td>We follow Antràs et al. (2017). They find a median markup of 35 percent across establishments. This estimate implies an elasticity of substitution of $\sigma = 3.85$.</td></tr>
<tr><td>Consumption share of manufactured goods (α)</td><td>We construct this parameter based on estimates by Duarte (2020), who uses detailed data on household consumption expenditure from the International Comparisons Programs (ICP) to compute consumption expenditures and relative prices of manufactured goods and services in many countries. She computes a real share of manufactured goods consumption in all consumption expenditures of $45\% - 50\%$ for high-income countries such as the U.S. (Table 4).</td></tr>
<tr><td>Dispersion of productivities (ζ)</td><td>We set this parameter based on Eaton and Kortum (2002), who estimate it from a gravity equation that relates bilateral trade flows to the characteristics of the trading partners and the distance between them.</td></tr>
<tr><td>Wages (w_n)</td><td>We estimate wages as two thirds times GDP divided by the size of the labor force (i.e., GDP per worker) from the World Bank World Development Indicators (WDI) in 2016. For each country we obtain GDP in current USD (series NY.GDP.MKTP.CD) and the total size of the labor force (series SL.TLF.TOTL.IN). For RoW, we take anaverage across the US' top-ten trading partners (listed in Table 2) using US imports from each country in 2016 as weight. US is normalized to 1.</td></tr>
<tr><td>Labor force (L_n)</td><td>From the World Development Indicators (WDI) in 2016 (series SL.TLF.TOTL.IN). For RoW, we sum the labor force of the top ten US trading partners in the period 1992-2016. US is normalized to 1.</td></tr>
<tr><td>Rate of trade wars U.S.-China ($\rho_{US,CN}$)</td><td>We take all J buyer-seller ($mxhcz$) quintuples in our data, identified as those where the associated $mhcz$ quadruple is in the first quartile of the within-country-product-mode (hcz) SPS distribution in the entire dataset. We compute for these the probability that a relationship separates after τ quarters, separately for China and RoW $$S_{c\tau} = \frac{\sum_{mxhzt} \mathbb{I}^T(\tau_{mxhczt} = \tau)}{\sum_{mxhzt} \mathbb{I}(\tau_{mxhczt} = \tau)}$$ where $\mathbb{I}(\tau_{mxhczt} = \tau)$ is equal to one if quintuple $mxhcz$ is at age $\tau_{mxhczt} = \tau$ quarters in quarter t, and $\mathbb{I}^T(\tau_{mxhczt} = \tau)$ is equal to one for all such quintuples that additionally trade for the last time in quarter t. We then fit the exponential decay function $e^{-\psi_{US,i} t}$ to the estimated separation probabilities to minimize the squared deviation for i =China and i =RoW. Since many quintuples trade only once, we fit this function from quarter two onwards, $\tau = 2, ..., 100$. We obtain $\psi_{US,RoW} = 0.0873$ and $\psi_{US,CN} = 0.1137$ yielding a difference of $\rho_{US,CN} = 0.0264$.</td></tr>
</table>

I Additional Estimation Details and Robustness

I.1 Baseline Estimation

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

$$\arg\min_{\phi \in \mathbb{F}} \sum_x T(\mathcal{M}_x(\phi), \hat{\mathcal{M}}_x) \tag{A.8}$$

where $T(\cdot)$ is the percentage difference between the model, $(\mathcal{M}_x(\phi))$, and data, $(\hat{\mathcal{M}}_x)$, moments, and $\mathbb{F}$ is the set of admissible parameter vectors, which is bounded to be

Table A.16: Construction of Empirical Moments

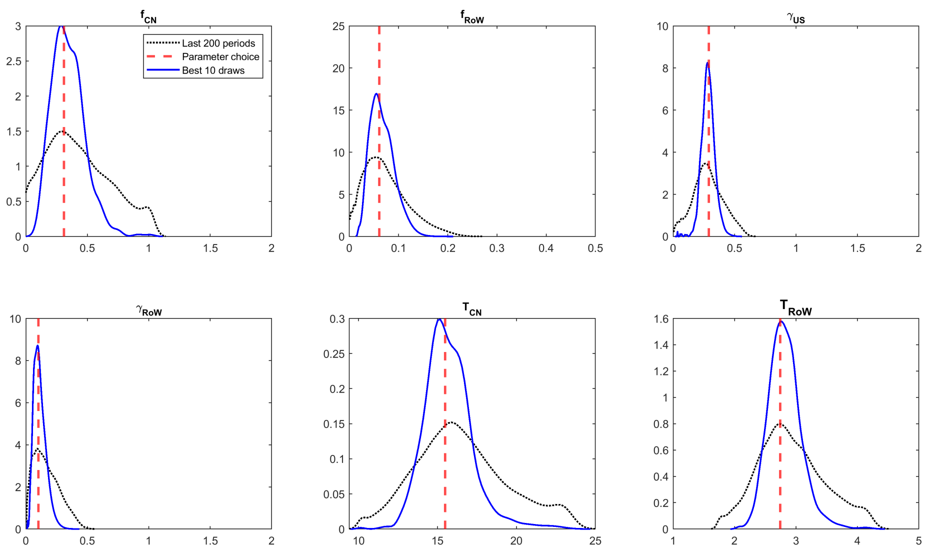
Moment	Description
Share of Chinese imports in domestic manufacturing sales	We target the US import penetration from China in 2016, computed as $IP_{CN} = \frac{\text{Imports}_{CN}}{\text{Domestic output} + \text{Total imports} - \text{Total exports}}$, where Imports_{CN} are US imports from China from https://www.census.gov/foreign-trade/balance/c5700.html, Domestic output denotes gross output in the manufacturing sector from https://www.bea.gov/itable/gdp-by-industry, and Total imports and Total exports are U.S. imports and exports of goods from https://www.census.gov/foreign-trade/balance/country.xlsx
Share of rest of world imports in domestic manufacturing sales	We target the US import penetration from the rest of the world in 2016, computed as: $IP_{RoW} = \frac{\text{Imports}_{RoW}}{\text{Domestic output} + \text{Total imports} - \text{Total exports}}$ where Imports_{RoW} are US imports from all countries except China from https://www.census.gov/foreign-trade/balance/country.xlsx.
Standard deviation of $\hat{\epsilon}$	We take the residuals from (14) and retain only those that have $\overline{WBS}_{mhcz}$ in the fourth quartile of the WBS distribution, i.e., those most likely associated with A sourcing, separately for imports from China and from the rest of the world. We collapse the residuals to the HS10 level to remove variation in shipping frequency within the same product that is unrelated to inspection costs and then take the standard deviation of the resulting product-level average residuals.

strictly positive and finite. In the choice of the function $T((\mathcal{M}_x(\phi), (\hat{\mathcal{M}}_x))$ we follow Lise et al. (2016) 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 Lise et al. (2016) and Engbom and Moser (2022), adapted to our needs. We simulate, using Markov Chain Monte Carlo for classical estimators as introduced in Chernozhukov and Hong (2003), 100 strings of length 1,000 (+ 200 initial scratch periods used only to calculate posterior variances) starting from 100 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 Gelman et al. (2013). The average parameter values across the 20 strings with the lowest values of the objective function provide a first estimate of the vector of parameters. We then repeat the same MCMC procedure, but we start each of our 100 strings from these parameter estimates.

Figure A.1 illustrates our approach. The black dotted line shows the density

Figure A.1: Estimation Outcomes

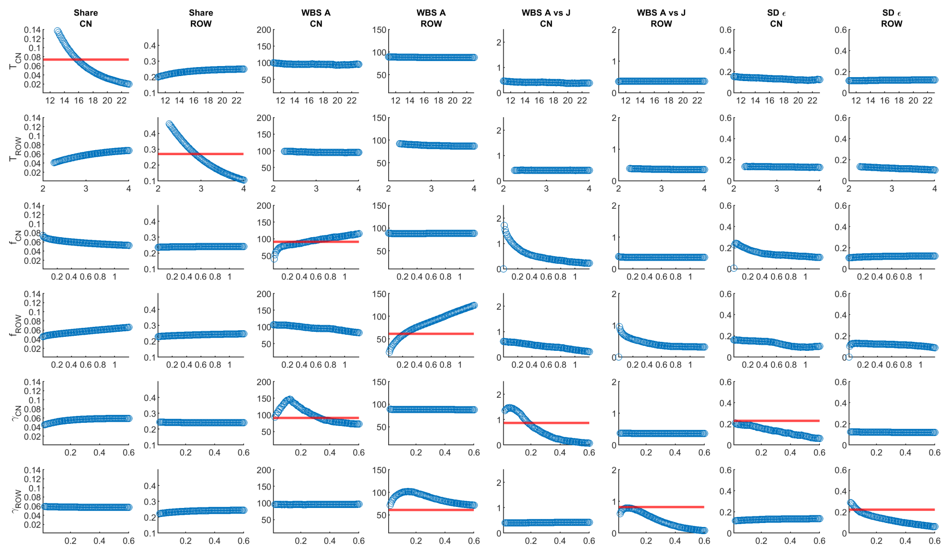


Source: Author's calculations, based on the estimation procedure described. Each panel shows the estimated parameter values for the parameter indicated in the title, under the assumption of a Pareto distribution for inspection costs. The black dotted line shows the density function of the parameter values associated with the last 200 iterations of our 100 strings. The red dashed line shows the average parameter values across the 100 best outcomes from all the draws. The blue density functions shows the density of the 10 best outcomes of each string, computed across all strings.

function of the parameter values associated with the last 200 iterations of our 100 strings. We pick the optimal parameters (red dashed lines) following Engbom and Moser (2022) as the average across the 100 best outcomes across all the draws. These correspond to the estimates reported in Table 10. For comparison, the blue density functions shows the density of the 10 best outcomes of each string, computed across all strings. This density provides an alternative way to select the best parameter values. All the densities are single-peaked, which suggests that the model is, at least locally, identified. Moreover, our chosen parameter values are generally very close to the peak of the densities.

Figure A.2 provides more detail on how each parameter is identified. We start from the optimal parameter values (red dashed lines in the previous figure) and vary each of the six parameters one-by-one on a grid of 100 values. For each parameter combination we solve the model 100 times, re-drawing the random productivity and inspection costs, and compute the average value of each moment. The panels in Figure A.2 plot the different values of each parameter (rows) against the values of the eight moments (columns). The main moments identifying the parameters are along the

Figure A.2: Identification of Parameters



Source: Author's calculations, based on the estimation procedure described. Each panel plots different values of the parameter indicated on the row against the moment indicated on the column, keeping all other parameters fixed at their optimal value. The blue dots show the averaged moment value across 100 runs with the given parameter choice, where the averaging is needed since the inspection cost and productivity draws differ across runs. The red horizontal lines represent the value of the moment in the data. We add these only for the main panels used to identify a given parameter in the data.

diagonal. The red horizontal line represents the value of the moment in the data, and hence identifies the parameter value that would lead the model to perfectly match this moment. While the relationships between the first four parameters and their main identifying moments are monotonic, for the last two parameters (the dispersion of inspection costs, γ_n) the relationships with some of the targeted moments are hump-shaped. Thus, there could be multiple values for each of these parameters that match a given moment equally well. We therefore target two sets of moments for these parameters (in the last four columns). This strategy yields a unique value for these parameters that minimizes the objective function. In Supplementary Appendix O.1, we perform an alternative exercise and plot the relationships between parameters and moments when we vary all parameters jointly. We show that the results are similar in this exercise.

Overall, these exercises highlight that our parameters of interest are well-identified from the moments we target.

I.2 Fréchet Distribution of Inspection Costs

We re-estimate the model using a Fréchet distribution instead of a Pareto distribution for the inspection costs:

$$G_n(m) = e^{-m^{-\gamma_n}}, \tag{A.9}$$

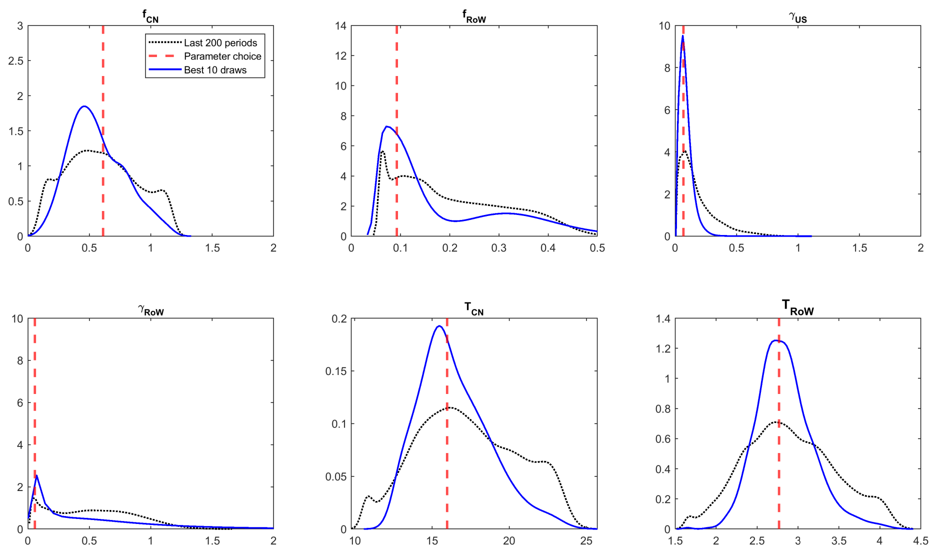
where γ_n is to be estimated. The other model parameters are set as before.

Figure A.3 presents our estimated parameter values analogously to Figure A.1. We find that all the densities are less tightly estimated than in the Pareto case. Our chosen parameter values are close to the peak of the densities.

Table A.17 presents the estimated parameter values and the values of the targeted moments in the simulations and in the data. The moments are reasonably well-matched, though less well than with the Pareto distribution. The model generates shares of Chinese and RoW imports in US manufacturing consumption that are close to the data, and generates shipping frequencies somewhat in line with their empirical analogues. The model does not match well is the difference in shipping frequencies between the first and the fourth quartile for shipments from China in row (5). In the data, the difference in shipping times between the first and the fourth quartile of the WBS_{mhcz} distribution is relatively small, while the dispersion of shipping times within the first quartile is relatively large. To match the latter the model estimates a high volatility of inspection costs (low γ_{CN}), which causes the model to overshoot the former moment for China. For the rest of the world, the two moments are relatively well matched. Due to this deviation from the targeted moments, we prefer the Pareto distribution as our baseline, which matches all moments better due to its different shape.

Table A.18 shows selected moments from our baseline equilibrium and the counterfactual without J relationships. Compared to the equilibrium with a Pareto distribution, the estimated share of J relationships is significantly higher for both China and for the rest of the world, with more than half of imports estimated to be under the J system. This higher share of J relationships results from the higher dispersion of inspection costs in this estimation, which generates more high inspection cost draws, leading J sourcing to be cheaper than A sourcing for more products. The structurally estimated J shares are in the ballpark of the empirical estimates we obtained using shipments in the first quartile of the SPS_{mhcz} distribution in Table 2. As a result of the higher share of J relationships, the welfare losses from removing such relationships rise by almost two percentage points compared to the baseline to 3.5 percent. The cost from eliminating J sourcing in the Fréchet case is therefore about two thirds as high as placing the US in autarky. This exercise suggests that the welfare losses from policy uncertainty can be much higher when the share of J relationships is greater.

Figure A.3: Estimation Outcomes with Fréchet Distribution



Source: Authors' calculations, based on the estimation procedure described, using a Fréchet distribution for inspection costs. Each panel shows the estimated parameter values for the parameter indicated in the title. The black dotted line shows the density function of the parameter values associated with the last 200 iterations of our 100 strings. The red dashed line shows the average parameter values across the 100 best outcomes from all the draws. The blue density functions shows the density of the 10 best outcomes of each string, computed across all strings.

Table A.17: Estimated Parameters and Targeted Moments

	(1)	(2)	(3)	(4)	(5)
	Parameter	Estimated Value	Moment that Primarily Identifies the Parameter	Moment in Data	Moment in Model
(1)	Productivity China (T_{CN})	15.973	Share of Chinese imports in domestic sales	0.074	0.049
(2)	Productivity RoW (T_{RoW})	2.769	Share of RoW imports in domestic sales	0.270	0.273
(3)	Fixed costs, China (f_{CN})	0.613	$\exp(\hat{\beta}_0 + \hat{\beta}_1 + \hat{\beta}_3\overline{beg} + \hat{\beta}_4\overline{end})$ from (14) for CN	91.00	105.49
(4)	Fixed costs, RoW (f_{RoW})	0.092	$\exp(\hat{\beta}_0 + \hat{\beta}_1 + \hat{\beta}_3\overline{beg} + \hat{\beta}_4\overline{end})$ from (14) for RoW	60.90	66.35
(5)	Dispersion of inspection	0.068	$\hat{\beta}_1$ from (14) for China	0.871	1.411
(6)	costs, China (γ_{CN})		Sd of $\hat{\epsilon}$ from (14) for China	0.227	0.187
(7)	Dispersion of inspection	0.056	$\hat{\beta}_1$ from (14) for RoW	0.822	0.726
(8)	costs, RoW (γ_{RoW})		Sd of $\hat{\epsilon}$ from (14) for RoW	0.219	0.238
(9)	Total objective $T(\cdot)$				0.580

Source: LFTTD and authors' calculations. Column (1) lists the parameters estimated for the model. Column (2) contains the estimated parameter values. Column (3) reports the moment targeted to identify the parameter. Column (4) presents the value of the moment in the data, and Column (5) presents the value of the moment computed in our simulated model.

Table A.18: Comparison of Equilibria with Fréchet Distribution

		(1)	(2)	(3)	(4)
		Baseline Equilibrium	**Equilibrium Without Japanese Sourcing**	**Autarky**	**Removal of PNTR**
(1)	Value imported from China (%)	4.9%	2.9%	.	4.5%
(2)	- of which, "Japanese"	56.7%	.	.	50.3%
(3)	Value imported from ROW (%)	27.3%	13.7%	.	27.4%
(4)	- of which, "Japanese"	67.6%	.	.	67.6%
(5)	Value imported from US (%)	67.8%	83.4%	100.0%	68.1%
(6)	Avg. inspection costs	0.2%	1.1%	.	0.2%
(7)	Avg. fixed costs (imports)	6.8%	4.4%	.	6.7%
(8)	Manufacturing price index	1.000	1.060	1.115	1.002
(9)	Utility	1.000	0.965	0.941	0.9994

Table shows various statistics of the equilibrium under the assumption of a Fréchet distribution for inspection costs. The first column presents the statistics for the baseline equilibrium, using the parameters that minimize the objective function. The second column shows the same statistics for a counterfactual economy in which the formation of "Japanese" relationships is not possible due to $\rho \to \infty$. The third column shows an autarky economy in which trade is not possible. The fourth column shows a counterfactual economy in which we reduce the arrival rate of trade wars from China to zero. Rows 1-5 show the share of US manufacturing sales, $P_{US}Q_{US}$, that is from China, from the rest of the world, and from the US , respectively, and the share of these manufacturing sales that is sourced under the "Japanese" system. Row 6 presents the average inspection costs as a share of the import value, computed over all imports, including under the "Japanese" system. Row 7 shows the average fixed costs as a share of the import value. Row 8 shows the manufacturing price index, P_{US}, normalized to one in the baseline. Row 9 shows total utility, $W_{US} = Q_{US}^{\alpha} Z_{US}^{1-\alpha}$, normalized to one in the baseline.

Supplemental Appendix

J Additional Proofs

J.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

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

Consider the case when $f+m=0$. If the condition holds for this case, it must also hold for $f+m>0$,because (S.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 (S.1) is equal to zero. Taking the derivative of the left-hand side of equation (S.1) with respect to y we obtain $1-e^{-y}(1-y)>0$. Thus, the left-hand side of (S.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 is positive for $f > 0$. Assume $f = 0$, and factor the numerator of $AC_J''(x)$ to obtain

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

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 $AC_J''(x) > 0$.

J.2 Continued Proof of Lemma 6.2: Average cost curves are convex and reach a limit

Part 1: Average cost curves are convex

American System Using (A.3) in Appendix A, the second derivative of average costs is

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

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

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

This expression holds if

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

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

Japanese System Equation (A.4) in Appendix A gives the slope of the average cost curve in the "Japanese" system. By the same arguments as in the "American" system $AC''(q) > 0$.

Part 2: Average cost curves reach a limit

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

$$q = \frac{\left[f + mw_b\right] re^{-r\psi_A}}{\bar{\theta}\frac{w_z}{\Upsilon}\left[1 - e^{-r\psi_A}\left[1 + r\psi_A\right]\right]}. \tag{S.2}$$

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

$$q = \frac{fre^{-r\psi_J}}{\bar{\theta}\frac{w_z}{\Upsilon}e^{(r+\rho)\psi_J}\left[(r+\rho)\psi_J\left[1-e^{-r\psi_J}\right]+1-e^{-r\psi_J}\left[1+r\psi_J\right]\right]}. \tag{S.3}$$

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.

K Additional Summary Statistics

We compare our baseline sample to an alternative arm's-length sample that does not restrict to buyer quadruples with at least five transactions. Since we cannot compute some variables such as weeks between shipments (WBS_{mhcz}) for quadruples that trade only a single time, we focus for consistency on the arm's length sample consisting of quadruples with two or more transactions.

Table S.1 presents an overview of the samples. The first column repeats some statistics of our baseline sample from Table A.1 in Appendix B. The second column presents the same statistics for the larger sample of quadruples with at least two transactions. The first row shows that the baseline sample accounts for slightly more than 80 percent of the broader sample of arm's-length trade by quadruples with at least two transactions. The next row shows that the broader sample contains almost twice as many importers, suggesting that most of the additional importers in the

Table S.1: U.S. Import Transaction Summary Statistics

	Baseline $t \geq 5$	Sample $t \geq 2$
Total Imports ($\$Bill$)	5,680	6,990
Unique Importers (m)	360,000	637,000
Unique Exporters (x)	5,037,000	6,531,000
Unique Importer-Product-Country-Mode Quadruples ($mhcz$)	2,966,000	7,615,000
Unigue Exporter-Importer-Product-Country-Mode Quintuples ($mxchz$)	21,700,000	30,600,000

Source: LFTTD and authors' calculations. Table summarizes U.S. arm's-length imports from 1992 to 2016. Observations are based on the cleaned data described in Appendix B. The first column restricts to our baseline sample of quadruples with at least five transactions ($t \geq 5$), analogous to Table A.1. The final column restricts to the broader sample of quadruples with two or more transactions ($t \geq 2$). Import values are in billions of real 2009 dollars. The final four rows of the table 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. Observation counts are rounded to the nearest thousand per U.S. Census Bureau disclosure guidelines.

broader sample do not have substantial imports. The third row presents the number of unique exporters and the fourth row shows the number of unique importer (m) by HS10 product (h) by country (c) by mode of transportation (z) $mhcz$ quadruples. The latter rises more than twofold in the broader sample. The last row presents the number of unique quintuplets. These do not increase nearly as much in percentage terms as the number of quadruples, as most of the quadruples unique to the broader sample have only few suppliers.

Table S.2 compares the $mhcz$ quadruples in the two samples. The first row shows that the average value traded by a quadruple in the broader sample is only about half of the trade value in the baseline sample. Rows two to four show that quadruples in the broader sample are shorter-lived, contain fewer shipments, and source from fewer suppliers on average. However, the average value per shipment is relatively similar to the baseline sample (row 5). Shipments in the broader sample are significantly more spaced out over time (row 6). The last two rows show that the average importer-exporter relationship length associated with a quadruple in the broader sample is shorter than in the baseline sample and that quadruples in the broader sample have a higher ratio of suppliers to shipments. The latter fact suggests that many of the additional quadruples not in the baseline sample conduct their few transactions with different suppliers.

Table S.3 shows statistics on the average number of sellers per shipment (SPS_{mhcz}) by main 6-digit NAICS industry of the importer, analogous to Table A.2 in Appendix B. For columns (3) and (4), we define J dummies J^k_{mhcz} that take a value of one if SPS_{mhcz} falls in the first quartile of its distribution within country-mode bins in the first time period ($k = cz$) to retain variation across products. Manufacturers are the most likely to use "Japanese" sourcing, consistent with these firms obtaining relatively customized inputs for their production processes.

Table S.2: Attributes of $mhcz$ Quadruples

	Baseline Sample $t \geq 5$		Broader Sample $t \geq 2$	
	Mean	*Standard Deviation*	*Mean*	*Standard Deviation*
Total Value Traded ($)	1,914,000	36,300,000	918,400	24,100,000
Length Between Buyer's First and Last Shipment (Weeks)	304.3	266	187.9	229.8
Total Shipments	38.6	157.9	17.8	100.4
Number of Sellers (x)	7.3	25.5	4.0	16.2
Value per Shipment (VPS), ($)	35,910	386,100	38,090	470,500
Weeks Between Shipments (WBS)	23.5	28.5	44.5	79.8
Average Relationship Length in Weeks ($length$)	180.8	154.7	147.2	156.7
Ratio of Sellers to Shipments (SPS)	0.334	0.241	0.512	0.306

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) quadruples during our 1992 to 2016 sample period. Observations are based on the cleaned data described in Appendix B. The first two columns restrict to our baseline sample of quadruples with at least five transactions, analogous to Table 1. The final two columns restrict to the broader sample of quadruples with two or more transactions. Observation counts are rounded to the nearest thousand per U.S. Census Bureau disclosure guidelines.

Table S.3: "Japanese" Relationships by Main Industry of the Importer

	Mean SPS		$J^{cz}_{mhcz} = 1$ Share of Import Value	
	(1)	(2)	(3)	(4)
Industry code (NAICS)	1995-2000	2002-2007	1995-2000	2002-2007
Manufacturing (31-33)	0.119	0.113	0.739	0.778
Agriculture (11)	0.123	0.106	0.584	0.630
Wholesale (42-43)	0.158	0.128	0.623	0.729
Other services	*0.160*	*0.130*	*0.655*	*0.713*
Professional services (54-55)	0.177	0.220	0.586	0.415
Mining, utilities and construction (21-23)	0.182	0.131	0.561	0.684
Finance and insurance (52-53)	0.187	0.213	0.516	0.514
Retail (44-45)	0.208	0.157	0.532	0.688
Information (51)	0.211	0.182	0.553	0.566
Admin support & waste mgmt (56)	0.213	0.195	0.312	0.423
Transportation and Warehousing (48-49)	0.216	0.210	0.487	0.511

Source: LFTTD and authors' calculations. The first two columns report the weighted average sellers per shipment (SPS_{mhcz}) across buyer quadruples with at least five transactions by main 6-digit NAICS industry-period. To obtain the main NAICS, we find in each year the industry with the importer's largest share of employment, and then take the modal main industry across the years in which the quadruple is active. We aggregate SPS_{mhcz} across quadruples using import values as weights. The second two columns report the share of the value of US imports accounted for by quadruples with SPS_{mhcz} in the first quartile of the distribution of SPS_{mhcz} within country-mode in the first period. Rows of the table are sorted by the column (1).

L Supplemental A vs J Classification Regressions

Differentiated Products Versus Commodities: We examine whether buyers are more likely to use J procurement for differentiated goods. If differentiated products have higher inspection costs, then by Proposition 2.1 buyers are more likely to use J procurement for them, which implies smaller shipment size, greater frequency, and higher unit import values than products sourced under the A system (Proposition 2.3). Moreover, as discussed in Section 3.3, this J 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,

$$\overline{Y}_{mhcz} = \beta_0 + \beta_1 Diff_h + \beta_2 \ln(VPW_{mhcz}) + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{cz} + \epsilon_{mhcz}. \tag{S.4}$$

We consider four dependent variables. The first is the average number of weeks between shipments WBS_{mhcz} as in the main text. 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.[58] Instead, we use as our second dependent variable the average transaction value per shipment, VPS_{mhcz}, as a measure of average transaction size. Third, we consider the average relationship length ($length_{mhcz}$) as in Section 3.3. Finally, the fourth variable is a measure of the buyer's procurement type, sellers per shipment (SPS_{mhcz}) introduced in the main text. 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).[59] 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 in the main text.

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

[59] 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 similar when we use the conservative definition.

The sample period is 1992 to 2016, we include only buyer quadruples with at least five transactions, and standard errors are clustered at the country-product level.

Results, reported in Table S.4, 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 J : they are shipped with fewer weeks between shipments, the average transaction size is smaller, and the average relationship length is longer. Results in the final column provide further support for this view, as buyer quadruples encompassing differentiated goods tend to have lower sellers per shipment.

Regressions by Sector: One concern with our findings could be that the results might only hold in some sectors, such as manufacturing, but not in others. We show in Tables S.5 to S.8 that our results regarding the relationship between SPS_{mhcz} and shipment attributes hold within different sectors: mining and utilities, manufacturing, wholesale, and retail.

A vs J Within Sellers: We next 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 J 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 J , but shipments within each seller might be dispersed if the buyer alternates among them. We use the following $mxhcz$-level OLS regression,

$$\begin{aligned} Y_{mxhcz} = \beta_0 + \beta_1 SPS_{mhcz} + \beta_2 \ln(QPW_{mxhcz}) + \beta_3 beg_{mxhcz} + \beta_4 end_{mxhcz} \\ + \lambda_{xhcz} + \epsilon_{mxhcz}. \end{aligned} \tag{S.5}$$

In this specification, Y_{mxhcz} represents procurement attributes at the buyer-seller relationship quintuple ($mxhcz$) 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 two-way clustered at the country (c) and product (h) level.

Results, reported in Table S.9, are similar to those in Section 3.2, providing further support for Proposition 2.3, as well as the use of SPS_{mhcz}. Across US buyer quadruples within foreign exporters, we find that increasing sellers per shipment by one standard deviation from its mean (from 0.33 to 0.58) is associated with a 5 log point rise in quantity per shipment, a 38 log point increase in weeks between shipments, a 3 log point decline in price, and a 16 log point drop in average relationship length.

Alternative Definition of Relationship Length: We next analyze the robustness of our measure of relationship length. If firms treat relationships with the same supplier across different products or modes of transportation as different relationships, then relationship length should not be defined using the time passed since the first ever transaction with the supplier overall but instead using the duration of the quintuple. We therefore construct an alternative relationship duration variable. First, for each $mxhcz$ quintuple, we compute the total number of weeks passed between the first and the last transaction. Second, for each $mhcz$ buyer quadruple, we take the average over the length of the $mxhcz$ quintuples within it. We refer to this variable as $Qlength_{mhcz}$ to indicate that it is based on the duration of the quintuple, rather than the overall length of the relationship between the importer and the exporter.

We run the same specification outlined in equation (7) using $Qlength_{mhcz}$ as the dependent variable. The results, reported in Table S.10, are similar to those in Table 4 in the main text, with coefficients that are about twice as large. The first column of the table shows that increasing sellers per shipment by one standard deviation from its mean is associated with a 61 log point decline in average relationship length. The second column shows that the average relationship length for quadruples in the fourth quartile is about 235 log points lower than the average relationship length for quadruples in the first quartile.

Table S.4: A vs J Classification Regression for Differentiated Goods

	(1)	(2)	(3)	(4)
Dep. var.	$\log(WBS_{mhcz})$	$\log(VPS_{mhcz})$	$\log(length_{mhcz})$	$\log(SPS_{mhcz})$
$Diff_h$	-0.234***	-0.225***	0.073**	-0.082***
	0.026	0.025	0.028	0.025
$\log(VPW_{mhcz})$	-0.464***	0.557***	-0.045***	-0.203***
	0.002	0.002	0.001	0.001
Observations	$2,589,000$	$2,589,000$	$2,589,000$	$2,589,000$
R-squared	0.611	0.730	0.193	0.278
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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country and product, are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table S.5: SPS_{mhcz} and Procurement Attributes - Mining and Utilities

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	0.413***	0.455***	-0.106**	-0.692***
	0.021	0.022	0.041	0.017
$\log(QPW_{mhcz})$	0.704***	-0.305***	-0.283***	-0.190***
	0.031	0.032	0.019	0.014
Observations	$25,500$	$25,500$	$25,500$	$25,500$
Fixed effects	hcz	hcz	hcz	hcz
R-squared	0.972	0.756	0.925	0.562
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 bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table S.6: SPS_{mhcz} and Procurement Attributes - Manufacturing

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	0.500***	0.538***	−0.181***	−0.540***
	0.014	0.014	0.022	0.012
$\log(QPW_{mhcz})$	0.769***	−0.238***	−0.367***	−0.131***
	0.018	0.018	0.022	0.008
Observations	560,000	560,000	560,000	560,000
Fixed effects	hcz	hcz	hcz	hcz
R-squared	0.950	0.712	0.816	0.434
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 bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table S.7: SPS_{mhcz} and Procurement Attributes - Wholesale

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	0.443***	0.475***	−0.181***	−0.571***
	0.015	0.015	0.013	0.020
$\log(QPW_{mhcz})$	0.682***	−0.328***	−0.281***	−0.167***
	0.012	0.012	0.017	0.007
Observations	1,215,000	1,215,000	1,215,000	1,215,000
Fixed effects	hcz	hcz	hcz	hcz
R-squared	0.945	0.708	0.856	0.469
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 bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table S.8: SPS_{mhcz} and Procurement Attributes - Retail

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	0.424***	0.458***	−0.120***	−0.556***
	0.030	0.031	0.023	0.022
$\log(QPW_{mhcz})$	0.643***	−0.366***	−0.195***	−0.115***
	0.007	0.007	0.012	0.008
Observations	525,000	525,000	525,000	525,000
Fixed effects	*hcz*	*hcz*	*hcz*	*hcz*
R-squared	0.945	0.708	0.856	0.955
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 bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table S.9: A vs J Classification Regression Across $mxhcz$ Quintuples

	(1)	(2)	(3)	(4)
Dep. var.	$\log(QPS_{mxhcz})$	$\ln(WBS_{mxhcz})$	$\ln(UV_{mxhcz})$	$\ln(length_{mxhcz})$
$\ln(SPS_{mhcz})$	0.100***	0.696***	−0.062***	−0.302***
	0.015	0.041	0.006	0.011
$\ln(QPW_{mxhcz})$	0.511***	−0.171***	−0.130***	−0.241***
	0.010	0.009	0.011	0.008
Observations	4,783,000	4,783,000	4,783,000	4,783,000
R-squared	0.966	0.621	0.953	0.786
Fixed effects	*xhcz*	*xhcz*	*xhcz*	*xhcz*
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 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}, P_{mxhcz}, and $length_{mxhcz}$ 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 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 quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) bins are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table S.10: SPS_{mhcz} and Alternative Relationship Length

	(1)	(2)
Dep. var.	$\log(Qlength_{mhcz})$	$\log(Qlength_{mhcz})$
$\log(SPS_{mhcz})$	-1.126^{***}	
	0.039	
$(SPS_{mhcz} = Q2)$		-0.653^{***}
		0.013
$(SPS_{mhcz} = Q3)$		-1.230^{***}
		0.024
$(SPS_{mhcz} = Q4)$		-2.348^{***}
		0.046
$\log(QPW_{mhcz})$	-0.164^{***}	-0.137^{***}
	0.008	0.006
Observations	$2,966,000$	$2,966,000$
R-squared	0.619	0.613
Fixed effects	hcz	hcz
Controls	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing the average quintuple relationship length within each quadruple ($Qlength_{mhcz}$) quadruples' sellers per shipment (SPS_{mhcz}), sellers per shipment quartile dummies 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 quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) bins are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

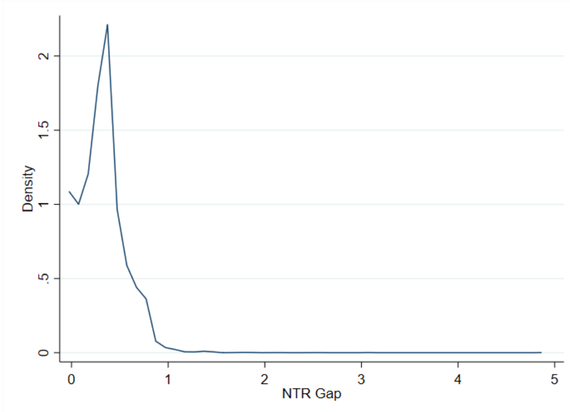
M Description of PNTR

This section provides more detail on the US granting permanent normal trade relations (PNTR) to China. US 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 US offers to trading partners that are members of the World Trade Organization (WTO). A provision of US 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, and this uncertainty continued throughout the 1990s. During this time, firms engaged in or considering US-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 US firms of the trade-dampening effects of this uncertainty in the 1990s, and Handley and Limão (2017) provide a theoretical basis for these effects that operates via suppressed entry by Chinese exporters.[60] Alessandria et al. (2024) show that uncertainty regarding the annual renewal of China's NTR status each summer reduced US imports from China, while also driving intra-year seasonal patterns in imports. When the United States 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. This plausibly exogenous policy change provides a useful opportunity for testing Proposition 2.1, i.e., whether a decrease in the probability of a trade war leads to the adoption of more "Japanese" sourcing.[61] Our strategy follows Pierce and Schott (2016) in defining a product's exposure to PNTR as the difference between

[60] Handley and Limão (2017) also estimate that the reduction in uncertainty associated with PNTR's ultimate approval was equivalent to a 13 percentage point permanent reduction in tariff rates.

[61] See also Blanchard et al. (2016), who examine how the presence of global value chains can affect the longer-term endogenous determination of tariff rates as part of multilateral trade negotiations.

Figure S.1: Distribution of the $NTR\ Gap$



Source: Feenstra et al., 2002 and authors' 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 US tariffs on Chinese goods might have risen absent the change in policy.

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,

$$NTR\,Gap_h = Non\ NTR\,Rate_h - NTR\,Rate_h. \tag{S.6}$$

We compute these gaps as of 1999, the year before the change in policy, using *ad valorem* equivalent tariff rates provided by Feenstra et al. (2002). As indicated in Figure S.1, these gaps vary widely across products, and have a mean and standard deviation of 0.32 and 0.23, respectively.

N Supplemental DID Regressions

mhcz Quadruple Level: In the main text we show that PNTR changed the shipping patterns (quantity per shipment, weeks between shipments, and unit value) at the $mxhcz$ level. We now examine whether the shift from A to J procurement in response to PNTR also altered the shipping patterns at the $mhcz$ quadruple level. Compared to the regressions of continuing relationships at the $mxhcz$ level, this regression aggregates across the supplier dimension, and computes shipping attributes of the quadruple using transactions with all suppliers. It also allows 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 $mhczt$-level DID regression,

$$\begin{aligned}\ln(Y_{mhczt}) =& \beta_1 1\{t = Post\} * 1\{c = China\} * NTRGap_h + \beta_2 ln(QPW)_{mhczt} + \\ & \beta_3 \chi_{mhczt} + \lambda_{mhcz} + \lambda_t + \epsilon_{mhczt}.\end{aligned} \tag{S.7}$$

As before, Y_{mhczt} represents one of the three procurement attributes: average quantity per shipment (QPS_{mhczt}), average weeks between shipments (WBS_{mhczt}), and average unit value (i.e. value divided by quantity) (UV_{mhczt}).

Results, displayed in Table S.11, show a significant decline in the average shipping size and weeks between shipments, consistent with a shift towards J procurement. 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 et al., 2020), including privately owned firms that tend to have lower prices than state-owned incumbents (Khandelwal et al., 2013). New suppliers might also charge low, introductory prices to gain market share, further dampening unit values.

All Relationships: We re-run our relationship-level PNTR regression (9) using both continuing and new relationships simultaneously for all buyer quadruples and sellers that appear in both. Specifically, we run a modified version of the regression,

$$\begin{aligned}\ln(Y_{mxhczt}) =& \beta_1 1\{t = Post\} * 1\{c = China\} * NTRGap_h + \beta_2 ln(QPW_{mxhczt}) + \\ & \beta_3 \chi_{mxhczt} + \lambda_{mhcz} + \lambda_x + \lambda_t + \epsilon_{mxhczt},\end{aligned} \tag{S.8}$$

where we use importer-product-country-mode of transportation ($mhcz$) fixed effects, exporter (x) fixed effects, and period (t) fixed effects. Our results in Table S.12 indicate that PNTR leads to a decline in the quantity per shipment and the number of weeks between shipments, and an increase in the unit value for this set of relationships, consisting with a shift to J procurement.

Table S.11: Within $mhcz$ Quadruple PNTR DID Regression

	(1)	(2)	(3)
Dep. var.	ln(QPS_{mhczt})	ln(WBS_{mhczt})	ln(UV_{mhczt})
$Post_t * China_c * NTRGap_h$	-0.043***	-0.058***	0.018
	0.014	0.013	0.024
$ln(QPW_{mhczt})$	0.436***	-0.584***	-0.207***
	0.018	0.018	0.026
Observations	738,000	738,000	738,000
R-squared	0.978	0.887	0.974
Fixed effects	$mhcz, t$	$mhcz, t$	$mhcz, t$
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_{mhczt}), (WBS_{mhczt}), and (UV_{mhczt}) are average quantity per shipment, average weeks between shipments, and average unit value (i.e. value divided by quantity) in period t. All regressions include $mhcz$ and period t fixed effects, control for the beginning and end week of the quadruple as well as all variables needed to identify the DID term of interest. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

Table S.12: Within $mxhcz$ Quintuple PNTR DID Regression Using All Relationships: 2002-2007 vs 1995-2000

	(1)	(2)	(3)
Dep. var.	ln(QPS_{mxhczt})	ln(WBS_{mxhczt})	ln(UV_{mxhczt})
$Post_t * China_c * NTRGap_h$	-0.131***	-0.115**	0.078***
	0.012	0.012	0.027
$ln(QPW_{mxhczt})$	0.407***	-0.597***	-0.130***
	0.013	0.012	0.018
Observations	4,023,000	4,023,000	4,023,000
R-squared	0.966	0.838	0.971
Fixed effects	$mhcz, x, t$	$mhcz, x, t$	$mhcz, x, t$
Controls	Yes	Yes	Yes

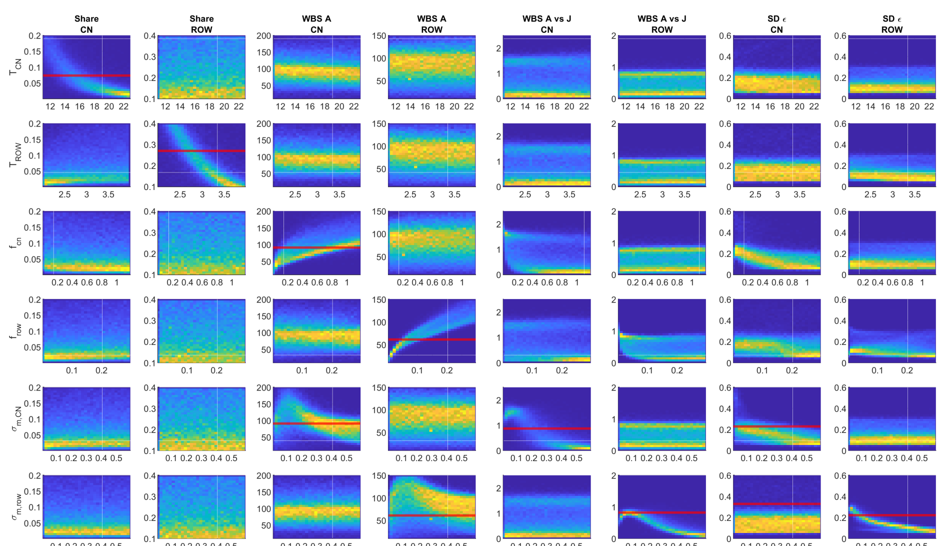
Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by exporter by product by country by mode of transport ($mxhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre-and post periods are 1995 to 2000 and 2002 to 2007. (QPS_{mxhczt}), (WBS_{mxhczt}), and (UV_{mxhczt}) are average quantity per shipment, average weeks between shipment, and average unit value (i.e. value divided by quantity) in period t. All regressions include $mhcz$, exporter x, and period t fixed effects, and control for the beginning and end week of the quadruple as well as all variables needed to identify the DID term of interest. Standard errors, adjusted for clustering by country (c) and product (h), are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

O Additional Quantitative Results

O.1 Identification

We perform an additional identification exercise. We vary all six parameters from the estimation jointly by drawing 100,000 different combinations of parameter values. We then simulate the model for each combination, obtain the simulated moments, and plot the resulting relationships between parameters and moments as a binscatter in Figure S.2. This exercise differs from Figure A.2, where we only varied one parameter at a time. The values of the six parameters are obtained as quasi random numbers drawn from a Sobol sequence. The figure shows similar relationships as Figure A.2, although the associations are noisier since all parameters vary jointly. In particular, there are strong and monotone relationships between the first four parameters and their targeted moments, and more hump-shaped relationships for the final two parameters.

Figure S.2: Joint Identification of Parameters



Source: Authors' calculations, based on the estimation procedure described. Each panel plots different values of the parameter indicated on the row against the moment indicated on the column, where all parameters vary jointly based on 100,000 random parameter draws from a Sobol sequence. Lighter colors indicate more frequently observed combinations of parameter values and moment values. The red horizontal lines represent the value of the moment in the data. We add these only for the main panels used to identify a given parameter in the data.

O.2 Additional Results

Figure S.3 provides further intuition for the welfare implications of eliminating J sourcing. The left and right panels display the share of expenditures of US imported versus domestically manufactured goods and welfare, respectively, as $\rho_{US,n}$ increases from zero to infinity.[62] As the trade war arrival rate rises, J sourcing declines as buyers switch to A sourcing for goods where the foreign productivity advantage is relatively large, and to domestic sourcing for goods where it is relatively low. These trade responses are most dramatic at initial increases in the arrival rate of trade war.

A source of welfare gains arising from changes in the arrival rate of trade wars is that J exports generate additional income due to the incentive premium (the second term on the right-hand side in equation (13)). For exports sold under the J system, the exporting country appropriates the incentive premium instead of having the foreign buyer country inspect the goods. As the arrival rate of trade wars rises, the number of products sourced under the J system falls. At the same time, a higher arrival rate of trade wars increases the incentive premium for each good that is still shipped under the J system.[63] The overall effect of these two opposing forces on US income, W_n, is highlighted by Figure S.4. There is an interior point which maximizes total US income, highlighting a potentially interesting avenue for trade policy. It is beneficial for a country to be associated with a lower arrival rate of trade wars, as this will allow its exporters to ship more under the J system and to collect the incentive premium. However, as the arrival rate of trade wars becomes too low, in our estimated equilibrium the reduction of the incentive premium dominates the extensive margin effect of additional products shipped under the J system. Thus, some trade policy uncertainty can be good to allow exporters to collect incentive premia.

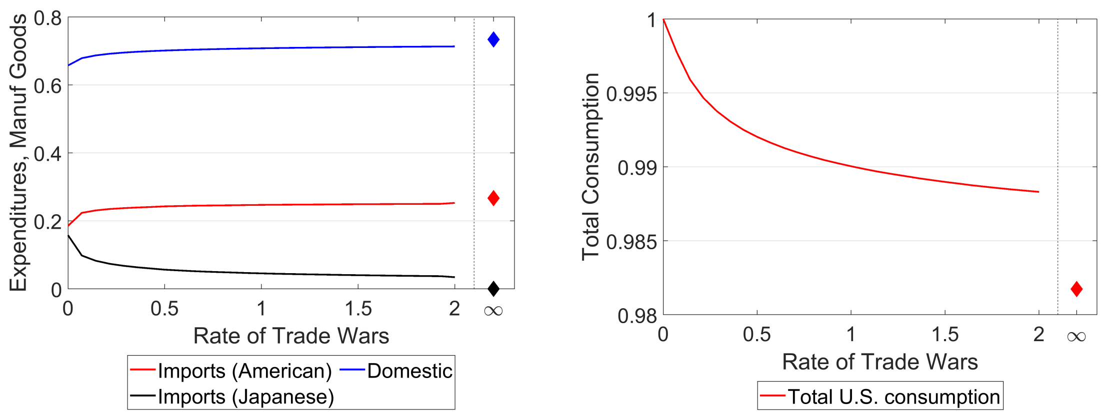
In our model the trade war arrival rate is symmetric for any country pair, and since importers always benefit from a lower arrival rate of trade wars overall welfare strictly falls with $\rho_{US,n}$, as shown in Figure S.3b. However, in a more general model in which $\rho_{n,i} \neq \rho_{i,n}$, a country would want to be perceived as slightly uncertain to maximize exporters' incentive income from J exports, while it would simultaneously want to commit its trade partners to never start a trade war to reduce import costs.

[62] We set the trade war arrival rate from China and from ROW to be equal in this exercise, $\rho_{US,CN} = \rho_{US,ROW}$, to facilitate the interpretation of the figure.

[63] Note from equation (2) that the incentive premium is positive even for $\rho = 0$.

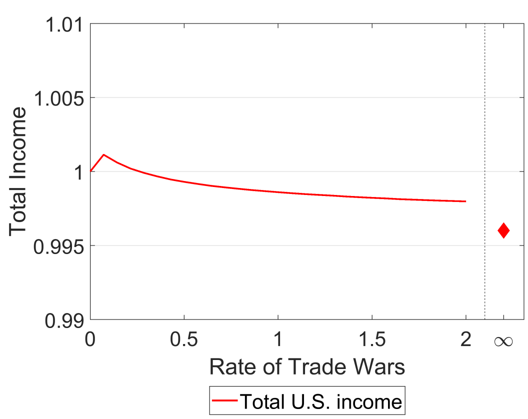
Figure S.3: Effect of Trade War Arrival Rate on Sourcing and Consumption

(a) Manufacturing Expenditures (b) Utility



Notes: The left panel displays the share of expenditures on manufactured goods by the United States as a function of the arrival rate of trade wars from the rest of the world, where we distinguish imports under the "American" system (red), imports under the "Japanese" system (black), and domestic sourcing (blue). The right panel shows US utility, calculated as $Q_{US}^{\alpha} Z_{US}^{1-\alpha}$, as a function of the trade war arrival rate from the rest of the world. Welfare at an arrival rate of zero is normalized to one.

Figure S.4: Effect of Trade War Arrival Rate on Income



Notes: The figure shows US total income, i.e., wage income plus incentive premia, normalized to one for the baseline case, as a function of the trade war arrival rate $\rho_{US,n}$.