Trade Policy Uncertainty and the Structure of Supply Chains*

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Abstract

We model the impact of changes in trade policy uncertainty on supply chains and show that a reduction in the probability of a trade war can foster the adoption of "Japanese"-style procurement practices, in which domestic buyers ensure the provision of high-quality inputs from foreign suppliers via long-term, just-in-time relationships. Empirically, we first show that the model provides a useful framework for analyzing shipments between U.S. importers and foreign exporters, and then demonstrate that a change in U.S. trade policy that eliminated the possibility of substantial increases in U.S. tariffs on Chinese goods coincides with a shift towards "Japanese" procurement. We estimate that the shift led to substantial U.S. welfare gains. (JEL F13, F14, F15, F23) (Keywords: Supply Chain, Uncertainty, Trade War, Procurement)

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1 Introduction

Motivated by the success of Japanese manufacturers such as Toyota, many firms around the world have introduced "Japanese"-style procurement practices in an effort to enhance operational efficiency.[1] A key feature of these systems – in addition to the much-studied just-in-time inventory management – is the presence of long-term relationships between buyers and sellers (Liker and Choi (2004)).[2] Given the increasingly global reach of supply chains (Baldwin and Lopez-Gonzalez (2015)), trade policy represents a potentially important – yet under-studied – consideration in the ability of buyers and sellers to establish such relationships. Indeed, if buyer and seller are located in different countries, a high probability of a trade war can inhibit foreign sellers from entering into the sort of long-term relationships with domestic buyers that characterize the "Japanese" system.[3] This disincentive can adversely affect firm performance and welfare in several ways. For example, trade policy uncertainty might raise buyers' costs by forcing firms to hold higher levels of inventory.

This paper examines the role of trade policy in firms' selection of procurement systems both theoretically and empirically. In the first part of the paper, we develop a model in which buyers face a trade-off between two stylized procurement systems defined in Taylor and Wiggins (1997). Under the "Japanese" system, buyers motivate sellers to maintain product quality by committing to long-run purchases at a price above sellers' costs. The opposing "American" system, by contrast, has buyers choosing the lowest-cost seller for each order via competitive bidding, and using costly inspection to deter cheaters from shipping low quality. We demonstrate that changes in sellers' beliefs about the probability of a trade war can induce firms to switch between the American and Japanese systems. In the second part of the paper we first show that our model captures key features of transaction-level U.S. import data, and then demonstrate that a change in U.S. trade policy that eliminated the possibility of substantial tariff increases on Chinese imports coincides with a relative shift towards Japanese-style procurement between U.S. buyers and Chinese sellers. In the final part

[1]This movement is documented in a series of studies. See, for example, O'Neal (1989), Heide and John (1990), Lyons, Krachenberg, and Henke Jr. (1990), Dyer and Ouchi (1993) , Han, Wilson, and Dant (1993), Helper and Sako (1995) and Liker and Choi (2004).

[2]More broadely, "Japanese"-style buyer-seller relationships are also characterized by joint learning and information sharing, though we do not examine these elements in this paper.

[3]A new but growing literature uses the detailed importer-exporter information in the U.S. trade data to observe the structure of supply chains and buyer-seller relationships, including Monarch (2015), Boehm, Flaaen, and Pandalai-Nayar (2015), Monarch and Schmidt-Eisenlohr (2015) and Heise (2015).

of the paper we combine the results of the empirical analysis with numerical simulations of the estimated model in general equilibrium to explore potential welfare gains associated with the change in U.S. policy.

Our theoretical analysis is based on the framework introduced by Taylor and Wiggins (1997), where buyers under the American system pay a fixed cost per shipment to inspect goods for quality, while buyers under the Japanese system incentivize sellers to provide high-quality goods through repeated payment of a price premium over the sellers' costs. Taylor and Wiggins (1997) demonstrate that shipments between seller and buyer are optimally smaller and more frequent – i.e., more "just-in-time" – under the Japanese system. Here, we embed the Taylor and Wiggins (1997) framework in an Eaton and Kortum (2002) style general equilibrium model of trade to extend the analysis to international procurement.

We demonstrate that the higher are sellers' (exogenous) beliefs about the probability of trade peace with another country, the more likely they are to enter into Japanese-style procurement relationships with buyers from that country. The intuition for this result is straightforward: the higher the belief about the probability of trade peace, the greater the seller's confidence that a long-term relationship with a particular buyer can be sustained. This increased confidence lengthens the time horizon over which the seller expects to collect a premium over their costs from exporting their intermediate good to the buyer, driving down the premium needed to incentivize quality and thereby the relative cost of the Japanese system compared to the American system.

In our empirical analysis, we examine some of the fundamental features of the Taylor and Wiggins (1997) model using transaction-level U.S. import data. Through the lens of the model, we classify importers as using either Japanese- or American-style procurement based on the number of foreign suppliers from which they purchase goods within a product-country bin. Purchases from many suppliers are interpreted as evidence of American-style procurement while purchases from a small number or even a single supplier are deemed evidence of a Japanese-style relationship. We then show that transactions classified as American exhibit larger, less frequent shipments at lower prices and overall shorter buyer-seller relationships, as implied by the model. We also demonstrate that products classified by Rauch (1999) as differentiated exhibit smaller, more frequent shipments and longer relationships than products which have reference prices or which are traded on organized exchanges. These findings accord well with the idea that differentiated products are more costly to inspect and are therefore more likely to be traded under the Japanese system. To our knowledge, these results provide the first systematic evidence supporting the key insights in Taylor and Wiggins (1997).

We then examine the core implication of our extended model, whether a change in U.S. trade policy that eliminated the possibility of a sudden spike in U.S. tariffs induced a relative shift towards Japanese procurement. This analysis exploits variation in the exposure of U.S. import products to the U.S. extension of Permanent Normal Trade Relations (PNTR) to China in October 2000. Following Pierce and Schott (2016), we measure the exposure of a product to this trade liberalization as the potential jump in the tariff rate that could have occurred before the change in policy. Our triple difference-in-differences specification asks whether U.S.-China transactions within importer-exporter-product bins change after the policy is implemented (first difference) relative to bins for other countries (second difference) in products with greater exposure (third difference).[4] In line with the model's predictions, we find that U.S.-Chinese shipments of more-exposed products become relatively smaller, relatively more frequent, and relatively higher priced – that is, more "Japanese"-style – after the change in policy. Coefficient estimates suggest that a one standard deviation increase in the *ex ante* potential jump in tariff rates is associated with a relative decline in average shipment quantity of 13 percent and an increase in average shipment price of 4 percent.

In the final part of the paper we present quantitative simulations of the model that incorporate changes in shipment patterns highlighted in our empirical analysis. These simulations reveal that the change in procurement patterns induced by the policy change increases U.S. imports from China by approximately 20 percent relative to previous levels, partly at the expense of other trading partners that were not subject to the policy change. The change in procurement patterns also has implications for U.S. welfare, which increases by 0.2 percent via a decline in final goods prices. This analysis suggests that changes in trade policy can have a meaningful impact on trade flows and welfare by inducing firms to re-optimize with respect to procurement, even in the absence of other forces such as tariff changes or wage and productivity differences that are commonly associated with welfare gains from trade.

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

This paper makes contributions to several fields. The model we develop is to our knowledge the first to link trade policy to the choice of procurement patterns, and provides an alternate perspective on the large literature examining contractual frictions in international trade.[5] Indeed, one solution to the problem of hold-up in the decision to outsource may be relationship formation (Kukharskyy and Pflüger (2010)), i.e., the sharing of long-term gains in a repeated game. Here, we examine how long-term, "Japanese" relationships can overcome frictions associated with guaranteeing the provision of high-quality inputs. One attractive feature of our approach is that it yields predictions regarding shipment patterns that can be tested using transaction-level trade data.[6]

More broadly, our paper contributes to research examining the behavior of importers (e.g., Blaum, Lelarge, and Peters (2015)), the implications of trade wars (e.g., Ossa (2014)), information frictions in international trade (e.g., Cristea (2011)), trade policy uncertainty (e.g., Handley and Limão (2013), Handley (2014)), importer-exporter relationships in international trade (Heise (2015), Monarch and Schmidt-Eisenlohr (2015)), and the impact of supply-chain disruptions on output (e.g., Boehm, Flaaen, and Pandalai-Nayar (2015)).

The remainder of this paper proceeds as follows. Section 2 outlines our theoretical model. Section 3 describes the data and presents our empirical analysis. Section 4 contains our quantitative simulations. Section 5 concludes. An online appendix contains additional results.

2 Theoretical Model

Incomplete contracts, information asymmetries and contract enforcement are common problems when domestic buyers procure products from foreign suppliers. Observed organizational forms and contract structures are the result of firms optimally structuring their supply chain and procurement systems.

Existing models in the international trade literature focus on the trade-offs asso-

[5] See, for example, the survey by Antràs and Helpman (2008). Procurement within countries is a subject of considerable research in the industrial organization literature. See, for example, Tadelis and Zettelmeyer (2015), Cicala (2015) and Bajari et al. (2014).

[6] Our model also contrasts with existing models of heterogeneous firms and trade, in which producers balance fixed and variable costs in determining whether to export or engage in foreign direct investment (e.g., Melitz 2003, Bustos 2011). Here, as in Taylor and Wiggins (1997), however, the fixed and variable costs are endogenous to firms' choice of a procurement system.

ciated with forming related party (within-firm) versus arm's-length outsourcing relationships to allocate property rights and solve hold-up problems (Antràs (2003, 2005); Antràs and Helpman (2008); Feenstra and Hanson (2005); Fisman and Wang (2010)). In the case of asymmetric information, where effort to successfully produce components by the foreign supplier is not perfectly observed, upstream integration reduces the costs of monitoring (Grossman and Helpman (2004); Spencer (2005)). Integration is therefore one solution to mitigate contract and information frictions in international trade.

In this paper we provide a perspective on how firms engaged in international trade can mitigate contractual and information frictions when vertical integration is not an option, perhaps due to legal barriers. Taylor and Wiggins (1997) – hereafter TW – show that in this setting firms can solve a quality control problem using one of two procurement strategies, which they label the "American" and "Japanese" systems.

Under the American system, buyers use competitive bidding to select the lowest-cost supplier for each shipment, and use the threat of inspection to deter provision of low quality goods. Under the Japanese system, buyers incentivize honesty by purchasing exclusively from a single seller and indefinitely paying this seller a premium over her fixed and variable costs. TW demonstrate that shipments under the American system are larger and less frequent than under the Japanese system for two reasons. First, the fixed costs associated with inspection under the American system encourage buyers to minimize the number of orders. Second, sellers under the Japanese system have an incentive to order more frequently as a way of minimizing the payoff to a deviating seller. TW show that the optimal procurement choice depends on the ratio of the seller's fixed cost of producing each shipment to the buyer's fixed cost of inspecting each shipment. Intuitively, the lower the ratio of these fixed costs, the cheaper the Japanese system and the more likely it is to be embraced.

We use TW's model as a starting point to study how changes in trade policy affect firms' choice of procurement systems. We extend TW's model by assuming that firms evaluate future rent streams not only based on their rate of time preference, but also the likelihood of a trade war, defined as the imposition in the buyer's country of a prohibitively high tariff on the sellers' output. In this setting, trade policy can influence sellers' decisions to enter into Japanese-style procurement relationships by affecting their beliefs about the probability of a trade war taking place.

After setting up the contracting problem between buyer firms and seller firms, we embed this structure into a general equilibrium model of international trade, where the buyer firms in the contracting problem act as intermediaries that sell their purchases to consumers. A key difference between the framework developed here and other models based on Eaton and Kortum (2002) is that cross-country income differences can arise from variation in the likelihood of a trade war across countries, which affect the intermediaries' choice of procurement system when dealing with foreign suppliers, in addition to differences in productivity. Countries that have a relatively higher likelihood of a trade war are less likely to form Japanese-style procurement relationships between intermediaries and foreign sellers since future rents are discounted more heavily, which shifts firms towards the American system and on average raises procurement costs. A change in trade policy that makes trade peace more likely reduces procurement costs by making the Japanese system relatively cheaper, allowing intermediaries to purchase at lower costs and thus increasing consumer welfare.

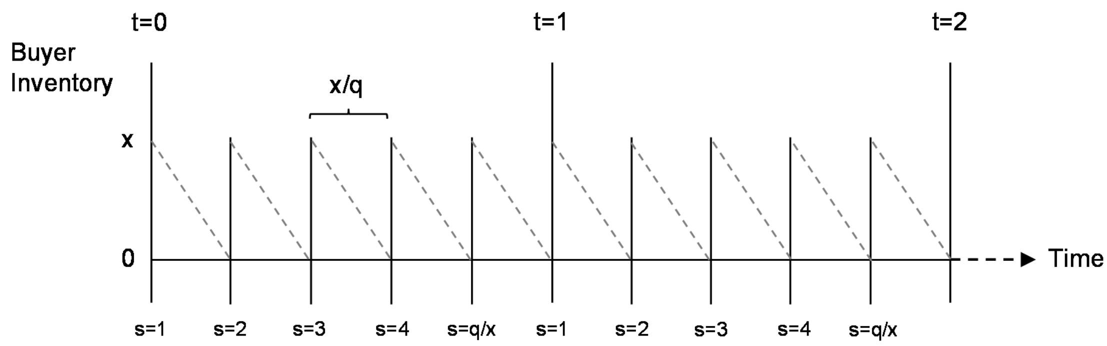
We develop this argument in this section. We first set up the contracting problem between the buyer and the seller firm, taking the quantity purchased by the buyer as exogenously given. After characterizing the model's solution, we embed it into a general equilibrium framework. We provide empirical support of the model in Section 3 and perform quantitative evaluations of the effects of a change in trade policy in Section 4.

2.1 Contracting Problem

Let time be denoted by t. In each period, a buyer seeks to purchase a fixed, exogenous quantity $q(t)$ from a seller firm, which may be located abroad. We will endogenize $q(t)$ below in the general equilibrium model. Since there will be no time-varying states in our model and therefore the quantities purchased $q(t)$ are the same in each period, we omit the time index from now on, and solve for an infinitely repeated static equilibrium.

The buyer firm receives quantity q in each period in a series of symmetric *shipments* of size x from the seller. As a result, there are q/x shipments during each time interval. The size and number of these shipments are chosen optimally by the buyer. Denoting the length of each period by Δt, these shipments arrive $\Delta t/(q/x)$ time intervals apart. We normalize $\Delta t = 1$, e.g., 1 year, so that the length of time passing between shipments is x/q. This shipment pattern is visualized in Figure 1.

We assume that the shipments incur costs at each point a shipment arrives. To finance these expenditures, intermediaries borrow their working capital requirements



Notes: Buyer uses inputs purchased from a seller at rate *q* over each time interval, e.g., *t*=0 to *t*=1. There are *s*={1,2,..*q/x*} shipments of size *x* per time interval, each arriving *x/q* time intervals apart. Between each shipment, the buyer's inventory falls from *x* to 0.

Figure 1: Model Setup

from (exogenous) banks at the beginning of the period. At the end of the period, the quantities q are shipped to households, households pay the firms, and the working capital loan is paid off.

We refer to the buyer-seller pair as a *sales relationship*. Sellers can choose to produce output of either low or high quality. Let $\theta \in \{\underline{\theta}, \bar{\theta}\}$ index the quality level of the output. Buyers require high quality, e.g., an acceptably low defect rate among the units shipped.[7] The sellers' problem is to determine whether to provide high- or low-quality goods for each shipment sent to the buyer.

Buyers can inspect each shipment at a cost m per shipment before accepting and paying for it. This cost captures, e.g., the complexity of the product shipped. Let α be the probability that such an inspection occurs. If a buyer chooses to inspect and the quality is low, the relationship with the current seller is terminated and the seller receives no payment from the buyer. We assume that goods are specific to the buyer, so that the seller cannot sell them to an alternative partner. Furthermore, if the seller ships low-quality goods and is found out her reputation is harmed and she is excluded from the market forever. If the buyer does not inspect, the order is accepted and the seller is paid. If the order subsequently turns out to be of low quality, the relationship is terminated. In that case, the buyer cannot recover payment from the seller but can substitute contemporaneous and future orders from an alternate seller. Here, too, a

[7]In an extension of their basic setup, TW consider the output market into which buyers sell and have buyers choose the optimal level of θ. They show that for a sufficiently small discount rate, the optimal level of quality demanded by buyers is arbitrarily close to the first best optimal level of quality.

seller found shipping low quality is excluded from the market forever.[8]

The Seller's Problem

The seller's production takes place according to the production function

$$x = \Upsilon l,$$

where Υ is a productivity term and l represents the labor input needed to produce quantity x. Workers are paid a wage w_z per unit of labor provided, which is paid after a batch is shipped. Here, w_z is the wage in the seller's country. We assume that the wage costs are proportional to the quality parameter θ, capturing the fact that producing high quality goods is more costly. Thus, sellers are faced with a total cost of $\theta w_z/\Upsilon$ per unit of output at each shipment, capturing labor costs and a quality premium. Sellers also incur a fixed cost f per batch shipped, which encompasses the fixed cost of both setting up and delivering a production run.[9]

The seller receives a payment of $v_s(x_s, \theta)$ per *shipment*, where s indicates whether the payment is under an American or Japanese system. We assume the seller does not have any bargaining power and fills an order only if she at least breaks even,

$$v_s(x_s, \theta) \geq f + \theta \frac{w_z}{\Upsilon} x_s. \tag{1}$$

Firms discount payments at the exogenously given per-period interest rate r. We assume free trade between the buyer's and seller's countries, but that a trade war (i.e., a prohibitive increase in the import tariffs as in Ossa (2014)) is possible. In the event of a trade war, the import tariff on the input rises enough to sever existing buyer-seller relationships between the affected countries.[10] Sellers' exogenous belief about the probability of continued peaceful trade between the buyer's and the seller's country, and therefore that the relationship will be maintained, is $e^{-\rho}$, where $\rho > 0$ reflects the rate at which trade war shocks arrive.[11] With continuous discounting, the discount

[8]This assumption is a simplification. In actuality, practitioners of Japanese procurement tend to reduce orders to suppliers that ship sub-standard goods but do not eliminate them unless violations are egregious or not corrected. See, for example, Liker and Choi (2004).

[9]Thus, we ignore any transportation costs which depend on shipment size or value. We note that uncertainty over these costs may also inhibit the formation of long-term relationships.

[10]Recent research (Ossa (2014)) indicates that the optimal tariffs countries might set in the event of a trade war are substantial, averaging 63 percent worldwide.

[11]The model considers trade in a single product. An alternate interpretation of ρ that brings the model closer to our data analysis below is that it reflects both the probability of a trade war (which

factor between shipments is then given by $e^{-(r+\rho)x_s/q}$.

Given that the stationary environment described above implies a continuous repetition of order cycles over time, the net present value to the seller of supplying shipments of x_s to the buyer as $T \to \infty$ is

$$\frac{v_s(x_s,\theta) - f - \theta\frac{w_z}{\Upsilon}x_s}{1 - e^{-(r+\rho)x_s/q}}. \tag{2}$$

As a result, the seller ships high quality ($\theta = \bar{\theta}$) if and only if expression (2) is at least as great as the one-time profit from cheating by supplying low quality ($\theta = \underline{\theta}$), i.e.,

$$\frac{v_s(x_s,\bar{\theta}) - f - \bar{\theta}\frac{w_z}{\Upsilon}x_s}{1 - e^{-(r+\rho)x_s/q}} \geq (1-\alpha)v_s(x_s,\bar{\theta}) - f - \underline{\theta}\frac{w_z}{\Upsilon}x_s. \tag{3}$$

As this expression makes clear, decreases in shipment size x_s, as well as decreases in the arrival rate of trade war shocks, ρ, raise the seller's discount factor, thereby strengthening the seller's incentive to provide high-quality shipments.[12]

The Buyer's Problem

The buyer chooses to conduct procurement either under the American (A) or the Japanese (J) system. Under the American system, buyers select the lowest cost supplier and use inspections to deter cheating. To simplify the problem we assume buyers under the American system always inspect while buyers in the Japanese system never inspect, so that $\alpha_A = 1$ and $\alpha_J = 0$.[13] Given our assumptions, under the American system, the seller just breaks even on each shipment,

$$v_A(x_A,\bar{\theta}) = f + \bar{\theta}\frac{w_z}{\Upsilon}x_A. \tag{4}$$

As there is no expectation of a long-term relationship under the American system, this shipment value satisfies the seller's incentive compatibility constraint (equation (3)).

Under the Japanese system, buyers obtain seller honesty through repeat purchases

is the same for all products) and the subsequent rise in tariffs (which might vary across products) for the particular good being traded. The probability of breakup is rising in the latter.

[12] An alternative approach to incorporating trade policy uncertainty would be to multiply the discount factor by an exogenous probability of trade peace $(1-\rho)$. However, a drawback of this approach is that then the probability of relationship separation over a given time period is not independent of the number of shipments made.

[13] TW show that optimal inspection under the American system is a function of shipment size and quality, $\alpha_A^* = \overline{\alpha}(x,\theta) > 0$, while under the Japanese system inspections do not occur, $\alpha_J^* = 0$.

and by paying sellers a premium over their costs. The shipment value under the Japanese system is

$$v_J(x_J,\bar{\theta}) = f + \bar{\theta}\frac{w_z}{\Upsilon}x_J + \left(\frac{1}{e^{-(r+\rho)x_J/q}} - 1\right)(\bar{\theta} - \underline{\theta})\frac{w_z}{\Upsilon}x_J. \tag{5}$$

This equation holds with strict equality given the assumption that the buyer holds all the bargaining power, but is still incentive compatible for the seller. The third term on the right hand side reflects the premium over the shipment value paid under the American system, $v_A(x_s,\bar{\theta})$, that a buyer under the Japanese system pays to incentivize the seller to sustain high quality over a long-term relationship. Intuitively, this premium rises as the rate at which trade wars arrive, ρ, rises.

Let $\Psi(x_s) = e^{-(\rho/q)x_s}$ be the per shipment probability of continued peaceful trade. We assume that in the event of a trade war, the relationship is terminated and buyers switch to an alternate supplier from a different country. We define the net present discounted cost of procurement from that alternate supplier by $\hat{C}$. The net present cost to the $s = \{A, J\}$ buyer of procuring a total quantity q using batch size x_s for each shipment is then (excluding the interest costs on the working capital loan):

$$C_s(x_s(q)) = \frac{\Psi(x_s(q))\left[v_s(x_s(q),\bar{\theta}) + \alpha_s m\right] + (1-\Psi(x_s(q)))\,\hat{C}}{1 - e^{-(r+\rho)x_s(q)/q}}, \tag{6}$$

where we write $x_s(q)$ to make explicit the dependence of the batch size of the total quantity demanded per period. The equation states that at each shipment, with probability $\Psi(x_s(q))$ no trade war has occurred since the last shipmenet and the buyer makes a payment of $v_s(x_s(q),\bar{\theta})$ plus, under the American system, the inspection cost. With complementary probability, the relationship ends forever and the buyer receives her outside option (cost) $\hat{C}$. The discount factor in the denominator takes into account the probability of a trade war since the payments are only made up to that point.

The outside option value $\hat{C}$ may be chosen to reflect a number of scenarios. If $\hat{C} = 0$, then a trade war forces the buyer to stop purchasing the product forever. Going forward, we assume that $\hat{C} = C_s(x_s(q))$, so that the buyer is able to replace the lost relationship with a substitute that allows her to continue purchases under the same shipment pattern. Given these assumptions, we have that

$$C_s(x_s(q)) = \frac{v_s(x_s(q),\bar{\theta}) + \alpha_s m}{1 - e^{-r x_s(q)/q}}, \tag{7}$$

where now discounting takes place at rate r since the buyer is always able to purchase from some source, subject to paying a cost under the Japanese system in the event of a trade war. We define $\delta(x_s) = e^{-rx_s/q}$.

Given a final good's demand q and a final good's price p, the intermediary's problem is to choose the optimal order size $(x_s)^*$ and the optimal procurement system $s = \{A, J\}$. The buyer's discounted expected real profits of procuring quantity q in batches of x_s using system s are

$$\pi_s(x_s(q)) = \tfrac{e^{-r}}{1-e^{-r}} pq - C_s(x_s(q)) - C_s(x_s(q))(1 - e^{-r}), \tag{8}$$

where revenues pq accrue at the end of each period when households make purchases, and are hence discounted by e^{-r}. The last term in equation (8) reflects the borrowing costs for the working capital. At the beginning of each period, buyers take up a loan of $C_s(x_s(q))(1 - e^{-r})$ to cover their expected working capital requirements for the period. Given interest costs of $e^r - 1$, buyers expect net present interest payments of $C_s(x_s(q))(1 - e^{-r})$ at the beginning of each period when choosing their procurement strategy.

Since in partial equilibrium final goods prices and the quantity demanded q are taken as given, maximizing profits is equivalent to minimizing the cost function. In Appendix A.1, we show that the cost function is strictly convex in x_s if r is small. Then, the tradeoff associated with choosing lower- versus higher-frequency procurement can be seen by setting the first order condition of the cost minimization problem to zero, yielding

$$\frac{v_s'(x_s, \bar{\theta})}{1 - \delta(x_s)} = \frac{-\delta'(x_s)}{(1 - \delta(x_s))^2} \left(v_s(x_s, \bar{\theta}) + \alpha_s m \right), \tag{9}$$

where

$$v_s'(x_s, \bar{\theta}) = \begin{cases} \bar{\theta} \dfrac{w}{\Upsilon} & \text{if } s = A \\ \bar{\theta} \dfrac{w}{\Upsilon} + \left(\bar{\theta} - \underline{\theta} \right) \dfrac{w}{\Upsilon} \left\{ \left[1 + \left(\dfrac{r + \rho}{q} \right) x_s \right] \frac{1}{e^{-(r+\rho)\frac{x_s}{q}}} - 1 \right\} & \text{if } s = J \end{cases} \tag{10}$$

The left hand side of equation (9) represents the discounted value of higher costs associated with a small increase in order size (i.e., a small decrease in order frequency). The right hand side measures the savings from an increased discount factor due to spacing orders further apart in time. Note that fixed order costs, f – a parameter of $v_s(x_s, \bar{\theta})$ – and m, appear only on the right hand side of the expression: the higher these costs, the greater the benefit of raising order size (i.e., a small decrease in order frequency).

2.2 Characterization

We now provide analytical results for the optimal shipment pattern and comparative statics of the model, and illustrate our findings numerically. Throughout the rest of this paper, we normalize $\underline{\theta} = 0$. We begin our analysis by proving that the optimal order size under the American system, x_A^*, is larger than the optimal order size in the first-best (FB) scenario, x_{FB}^*, where neither inspection nor payment premia are required to deter provision of low-quality goods.[14]. On the other hand, the optimal order size under the Japanese system, x_J^* is smaller than the optimal order size in the first-best case.

Proposition 1. *The optimal shipment size satisfies* $x_A^* > x_{FB}^* > x_J^*$.

Proof. See Appendix A.2. □

The intuition for this result is straightforward. Under the American system, the buyer encourages the seller to supply high quality by paying a fixed inspection cost at the arrival of each shipment. In order to economize on these costs, the buyer optimally orders less frequently, which leads to larger batch sizes. Under the Japanese system, on the other hand, the seller is incentivized to provide high quality via premia, which she collects at each shipment. Smaller, more frequent shipments lead to lower discounting of future rent streams, which improves the seller's incentives.

Using equation (7) at the optimal order quantity, we find after application of the envelope theorem that

$$\frac{\partial C_A(x_s(q))}{\partial \rho} = 0 \tag{11}$$

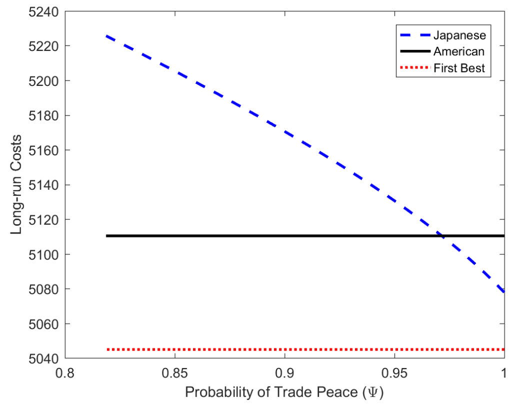
under the American system and

$$\frac{\partial C_J(x_s(q))}{\partial \rho} = \frac{\frac{x_J^*}{q} e^{(r+\rho)x_J^*/q} \bar{\theta} \frac{w_z}{\Upsilon} x_J^*}{1 - \delta(x_J^*)} > 0 \tag{12}$$

under the Japanese system. These equations highlight two key properties of our model. First, an increase in the likelihood of a trade war ρ (a decrease in the probability

[14]The quantity x_{FB}^* is the solution to the problem under the American system conditional on the fixed inspection cost (m) being zero.

Figure 2: Overall Cost vs Continuation Probability (Ψ)



of trade peace) does not affect costs under American procurement, since under that system incentives are provided via inspections and there are no switching costs. Second, a higher likelihood of a trade war raises total procurement costs under the Japanese system. When relationships are more likely to break up, it becomes harder for the buyer to provide incentives to the seller, forcing her to pay a higher premium over marginal costs.

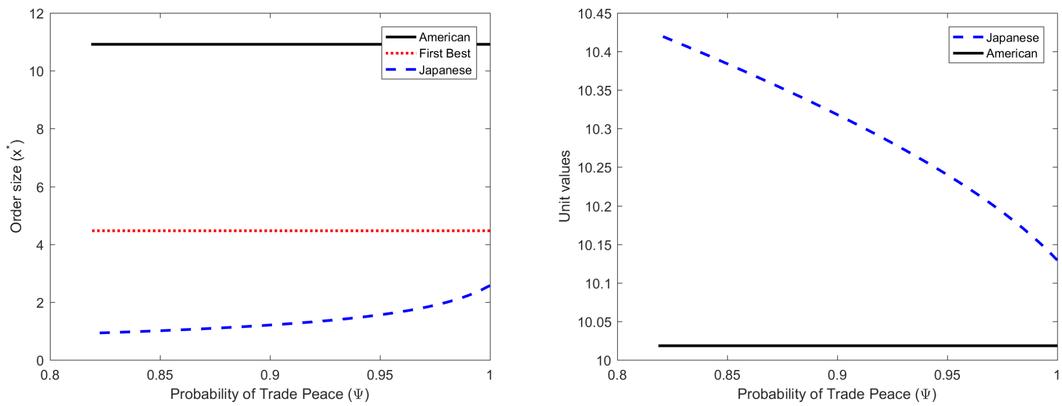
We illustrate this property of our model graphically in Figure 2. We impose the baseline parameters listed in Table A.1 in Appendix B.1 in this simulation. Define $\Psi \equiv e^{-\rho}$ as the per-period probability of trade peace. Figure 2 confirms that costs are unaffected by an increase in the probability of trade peace under the American system and in the first-best scenario, while they decline under the Japanese system.

Two other features of Figure 2 are worth noting. First, it shows that even if $\Psi = 1$ the cost of the Japanese system does not drop to that of the first-best scenario, The reason for this outcome is that even when trade peace is assured, the seller must be compensated for discounting if, as is the case here, $r > 0$. Second, Figure 2 reveals that beyond some threshold level for Ψ, which we denote Ψ^{Switch} (arbitrarily equal to 0.97 in the figure), the cost of the Japanese system drops below that of the American system. At that point, buyer and seller switch from the American to the Japanese system.

Proposition 2. *For a given set of parameters and $m \geq \underline{m} > 0$, where $\underline{m}$ is a threshold value, there exists $0 < \Psi^{Switch} \leq 1$ such that the American system is chosen for $\Psi < \Psi^{Switch}$ and the Japanese system is chosen for $\Psi > \Psi^{Switch}$. If $m < \underline{m}$ then the American system is always chosen.*

Figure 3: Order Size and Price vs Continuation Probability (Ψ)

(a) Order Size (b) Price



Proof. See Appendix A.3. □

To see the intuition for this proposition, note that for $m = 0$ the American system corresponds to the first-best solution and is therefore preferred to Japanese-style procurement since the payment of incentive premia can be avoided. However, since the procurement cost under the Japanese system is finite when $\Psi = 1$, and since procurement costs under the American system are independent of Ψ, there must be a threshold level $\underline{m}$ such that the Japanese system is preferred once the inspection cost exceeds this threshold level. As m is increased beyond $\underline{m}$, the Japanese system becomes preferred for additional values of $\Psi < 1$ and the switching point shifts to the left.

We now turn to how the optimal shipment size under the Japanese system changes with the probability of trade peace.

Proposition 3. *The optimal order size under the Japanese system, x_J^*, increases with the probability of trade peace Ψ.*

Proof. See Appendix A.4. □

We illustrate this property of the model in the left panel of Figure 3. When the probability of trade peace increases, it becomes easier for the buyer to provide incentives under the Japanese system, which enables her to economize on the shipping costs f by ordering less frequently. Under the American system and under the first-best scenario, the optimal shipment size does not depend on the probability of trade peace.

The price under the Japanese system, $\tau_J(x, \bar{\theta})/x_J^*$, falls when the probability of trade peace rises. The proof of this claim is straightforward. On the one hand, x_J^* is declining in the probability of trade peace, as shown in Proposition 3. On the other hand, $\partial C_J(q)/\partial\rho > 0$, as shown in equation (12). It must then be the case that $\partial\tau_J(x_J^*, \bar{\theta})/\partial\rho > 0$. This effect is shown in the right panel of Figure 3. The rise in order price as Ψ falls reflects the increase in the seller's rent necessary to incentivize her to produce high quality.

The key relationships for our empirical analysis come from joint consideration of Figures 2 and 3. Together, they reveal that if an increase in the probability of trade peace causes Ψ to jump from below Ψ^{Switch} to above this level, observed order size *falls* and observed order price *rises* as buyer and seller switch from the American to the Japanese system, i.e., from the solid black lines in the figure to the dashed blue lines. This implication of the model allows us to distinguish empirically between a change within a given procurement system and a switch of systems. The empirical results reported in Section 3 are consistent with PNTR leading to a switch to the Japanese system in U.S.-China procurement in 2001.

The comparative statics with respect to the other model parameters are summarized in the following proposition.

Proposition 4. *The optimal order size satisfies the following properties :*

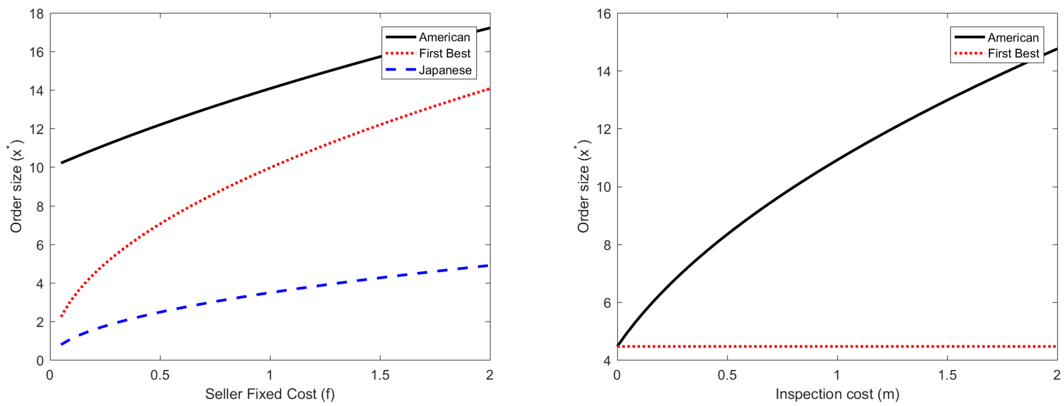
1. *Under both systems, the optimal order size x^* is increasing with the seller's fixed cost f*
2. *Under the American system, the optimal order size is increasing with the per-shipment inspection cost m*
3. *Under both systems, the optimal order size is decreasing in the marginal costs of high quality $\overline{\theta}$*
4. *Under both systems, the optimal order size x^* is increasing in total quantity demanded q. However, the time interval between shipments, x^*/q, declines with q.*

Proof. See Appendix A.5. □

The left panel of Figure 4 presents the relationship between order size and the seller's fixed cost f under both systems, while the right panel shows the relationship

Figure 4: Order Size vs Fixed Costs (f, m)

(a) Shipment Cost f (b) Inspection Cost m



between order size and per-shipment inspection cost m under the American system. In both cases, buyers seek to minimize incurring larger fixed costs by reducing shipments, thereby increasing order size.

The left panel of Figure 5 shows that optimal order size under both the American and Japanese systems declines with the marginal costs of high quality $(\overline{\theta})$. As the cost to produce high quality rises, buyers have an incentive to push purchases further into the future via more frequent, smaller orders. The right panel shows that the optimal order size increases with q under each system. An increase in quantity demanded causes the buyer to re-optimize by increasing the existing orders in order to fulfill the demand. However, the buyer also adjusts by ordering more frequently, as illustrated in Figure 6. When q goes up, the fixed order cost (and under the American system the inspection cost) become relatively less important since they can be spread over more units, which allows the buyer to order more frequently.

The effect of an increase in order size q affects total costs $C(x_s(q))$ differently under both systems. For r small, we can show that costs increase more strongly under the Japanese system than under the American system (See Appendix X). Intuitively, a larger quantity demanded diminishes the importance of the fixed inspection cost under the American system, while the incentive premium under the Japanese system is proportional to the individual order size, which rises with quantity. Thus, all else equal, our model predicts a shift towards the American system when overall trade increases. Hence, while a decrease in the probability of a trade war leads to the adoption of Japanese-style procurement practices, there is an offsetting force pushing

Figure 5: Order Size vs Variable Costs and Quantity $(\bar{\theta}, q)$

(a) Cost of High Quality $\bar{\theta}$ (b) Quantity q

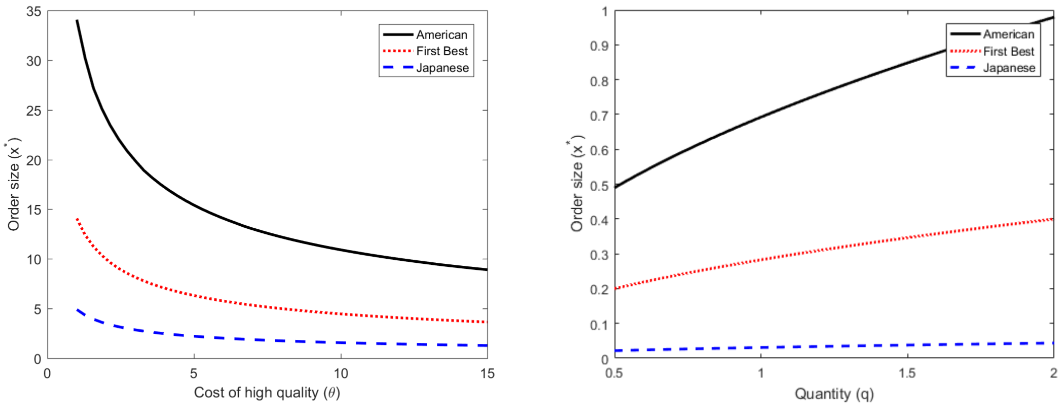
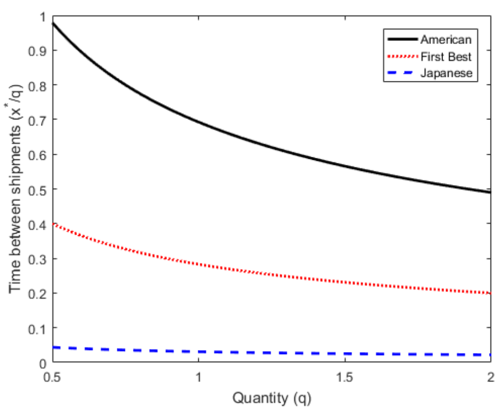


Figure 6: Time Between Shipments (x^*/q) vs Quantity (q)



in the other direction when the shift is accompanied by an increase in overall trade. Our empirical findings below suggest that the first effect dominates.

2.3 General Equilibrium

Households

We now embed the contracting problem into a general equilibrium framework in the spirit of Eaton and Kortum (2002). This setting will enable us to perform quantitative simulations of our framework and to assess the welfare consequence of a change in the probability of a trade war in Section 4.

Let there be a continuum of product varieties indexed by $j \in [0, 1]$. There is a finite number of countries N indexed by $n = 1, ..., N$. Each country n is populated by a set of households, which purchase a continuum of goods from buyer firms in their country. Going forward, we refer to these buyer firms as *intermediaries*, as they simply pass on purchases to consumers.[15] The quantity purchased of variety j by households in country n is denoted by q_n^j. Households aggregate the varieties according to a Dixit-Stiglitz aggregator

$$Q_n = \left[\int_0^1 (q_n^j)^{(\sigma-1)/\sigma} dj \right]^{\sigma/(\sigma-1)}, \tag{13}$$

where σ is the households' elasticity of substitution. Households provide labor in exchange for a nominal wage ω_n. Labor is completely immobile across countries and normalized to $L_n = 1$. We denote households' total labor income per period by W_n, and discuss its link with the wage below. Households also receive real profits π_n^j from intermediaries for each variety, and real profits $\hat{\pi}_n^j$ from manufacturers operating in their country. The households' objective is to maximize their total consumption Q_n subject to the budget constraint

$$\int_0^1 p_n^j q_n^j dj \leq W_n + P_n \int_0^1 \pi_n^j + P_n \sum_{n'} \int_{j \in \tilde{J}_{n'n}} \hat{\pi}_n^j, \tag{14}$$

where p_n^j is the final goods price of good j in country n, taken as given by households, P_n is country n's price index, and $\tilde{J}_{n'n}$ is the set of all varieties supplied by country n to country n'. A household's optimal consumption choice is thus

$$q_n^j = \left(\frac{p_n^j}{P_n} \right)^{-\sigma} Q_n, \tag{15}$$

and $P_n = \left[\int_0^1 (p_n^j)^{1-\sigma} dj \right]^{1/(1-\sigma)}$.

Sellers

There exists a sector of perfectly competitive sellers for each variety in each country, which can freely enter and exit the market. We indicate the seller country by i, and assume that Υ_i is a country-specific productivity parameter. In the simulations below, variation in these productivities across countries, in addition to differences in the

[15]It would be straightfroward to extend the model to allow the firms to transform the inputs.

probability of a trade war, motivates countries to trade. The real wage is defined by $w_i = \omega_i / P_i$, and hence the variable costs of a seller in country i are given by $\theta\omega_i/(P_i\Upsilon_i)$. We assume that the fixed costs depend both on the origin country i, the destination country n, and the variety j according to

$$f_{ni}^j = \psi^j(\omega_i/P_i) + d_{ni}^j,$$

where ψ^j is a product-specific cost associated with preparing the product for shipment, which is scaled by the wage level in country i, and d_{ni}^j is the average shipping cost of product j from origin country i to destination country n.

Intermediaries

We denote by $v_{ni,s}^j(x_{ni,s}^j(q_n^j), \bar{\theta})$ the payment of an intermediary in country n to a seller in country i for variety j under shipment system s, given a batch size of $x_{ni,s}^j$. The intermediary thus faces the cost function

$$C_{ni,s}^j(x_{ni,s}^j(q_n^j)) = \frac{v_{ni,s}^j(x_{ni,s}^j(q_n^j), \bar{\theta}) + \alpha_s m^j}{1 - e^{-r x_{ni,s}^j(q_n^j)/q_n^j}}, \tag{16}$$

where we assume that the inspection cost m^j is product-specific. Intermediaries are monopolistic competitors in their output market, choosing price to maximize their profits. For each variety j, seller country i, and shipment system s, intermediaries thus choose a desired consumer price p_n^j and a desired batch size $x_{ni,s}^j$ to solve

$$\pi_{ni,s}^j = \max_{p_n^j, x_{ni,s}^j} \left\{ \tfrac{e^{-r}}{1-e^{-r}} \frac{p_n^j}{P_n} q_n^j - C_{ni,s}^j(x_{ni,s}^j(q_n^j)) - C_{ni,s}^j(x_{ni,s}^j(q_n^j))(1 - e^{-r}) \right\}. \tag{17}$$

Intermediaries choose prices optimally taking into account households' demand curve. Setting the first-order condition of this problem with respect to price equal to zero yields

$$\frac{p_n^j}{P_n} = \frac{\sigma}{\sigma - 1} \frac{(2 - e^{-r})(1 - e^{-r})}{e^{-r}} ((C_{ni,s}^j)^*)'(q_n^j), \tag{18}$$

where we denote by $((C_{ni,s}^j)^*)'(q_n^j) \equiv (C_{ni,s}^j)'((x_{ni,s}^j)^*(q_n^j))$ the derivative of the minimized cost function at the quantity q_n^j. As usual, intermediaries charge a constant mark-up over marginal costs. However, in our problem, the intermediaries can affect these marginal costs by choosing the optimal shipping size. The choice of the optimal batch size is still determined by minimizing the cost function in equation (9).

Given the real profits $\pi^j_{ni,s}$ at the optimal batch size under each system, buyers choose the system $s \in \{A, J\}$ for each country which maximizes profits. Define $\pi^j_{ni} = \max\left\{\pi^j_{ni,A}, \pi^j_{ni,J}\right\}$ as country n's maximized profits of procuring from country i. Similar to Eaton and Kortum (2002), buyers in each country n order each good j from the country yielding the highest profits. Thus, the profits of an intermediary for variety j in country n are

$$\pi^j_n = \max\left\{\pi^j_{ni};\ i = 1, ..., N\right\}. \tag{19}$$

The actual batch size supplied will be

$$x^j_{ni,s} = \begin{cases} (x^j_{ni,s})^* & \text{if } \pi^j_n = \pi^j_{ni} \\ 0 & \text{if } \pi^j_n \neq \pi^j_{ni} \end{cases}.$$

These order sizes imply per-period shipping quantity q^j_n from country i to country n if country i is the low-cost producer, and a shipping quantity of zero otherwise.

A reduction in the probability of a trade war lowers the shipment cost for any order size under the Japanese system. Hence, for some varieties, there will be a switch to Japanese-style procurement. Our simulations below reveal that this change will lead to a fall in marginal costs, and hence in consumer prices. Note that the actual transactions price between the seller and the intermediary is higher under the Japanese system than under the American system, due to the incentive premium paid, but the overall procurement cost may be lower, as the Japanese system does not require the payment of the inspection cost. It is this saving that can be passed on to the consumer.

Equilibrium

Equilibrium in this economy requires that households maximize utility, firms maximize profits, and that the labor market clears. Labor market clearing implies for each country $i \in N$

$$1 = \sum_{n \in \tilde{N}_i} \int_{j \in \tilde{J}_{ni}} \frac{q^j_{ni}}{\Upsilon_i}, \tag{20}$$

where $\tilde{N}_i$ is the set of countries that country i exports to, $\tilde{J}_{ni}$ is the set of products exported to country n by country i, and q^j_{ni} is the quantity of good j purchased by country n from country i in each period. In equilibrium, each country purchases each good j from exactly one exporter, and thus $q^j_{ni} = q^j_n$ for $j \in \tilde{J}_{ni}$.

Households' budget sets at the beginning of the period must equal their actual income, which implies

$$W_i = \sum_{n \in \tilde{N}_i} \int_{j \in \tilde{J}_{ni}} \omega_i \frac{x^j_{ni,s}}{\Upsilon_i} \sum_{\tau=0}^{q^j_{ni}/x^j_{ni,s}-1} e^{r(1-\tau(x^j_{ni,s}/q^j_{ni}))}. \tag{21}$$

At each shipment to country n of good j, households earn labor income of $\omega_i(x^j_{ni,s}/\Upsilon_i)$, which is invested at interest rate r until the end of the period when purchases occur.

Definition 1. A competitive equilibrium consists of a vector

$$\left\{ \left\{ x^j_{ni,s} \right\}_{n,i,j,s}, \left\{ q^j_{ni} \right\}_{n,i,j}, \left\{ p^j_n, q^j_n \right\}_{n,j}, \left\{ \omega_n \right\}_n \right\}$$

such that

1. Given prices p^j_n and wages ω_n, households choose quantities q^j_n to maximize (13) subject to (14), for each n and j

2. Given quantities q^j_n , intermediaries set prices p^j_n and choose an order size $x^j_{ni,s}$ to maximize profits (17) for each country and system, choose a profit maximizing system s for each country, and a profit-maximizing country i (19), for each j, in each n

3. Wages ω_n are such that labor markets clear according to (20), in each n.

The model we propose resembles Eaton and Kortum (2002) in that each country chooses to procure each good from exactly one source country. However, while in Eaton and Kortum (2002) the choice of supplier is driven by cross-country productivity differences, we build a general equilibrium model in which not only productivity differences but also the probability of a trade war and the choice of optimal procurement system affect shipping patterns. As we show in the quantitative exercise in Section 4, changes in the probability of a trade war can give rise to switches in procurement strategies and optimal supplier choice which may entail significant welfare effects.

3 Empirical Analysis

We use transaction-level U.S. import data to examine the implications of our theoretical model. First, we develop a procedure for classifying U.S. importers as users of either the Japanese or American procurement systems and examine whether purchases by these firms differ along the dimensions suggested by the model. We then investigate whether transactions between U.S. buyers and Chinese sellers became more Japanese after a change in trade policy that increased the likelihood of trade peace between the United States and China.

3.1 Description of the Data

The U.S. Census Bureau's Longitudinal Foreign Trade Transaction Database (LFTTD) tracks every U.S. import transaction from 1992 to 2011. Data available include the dates the shipment left the exporting country and arrived in the United States, identifiers for the U.S. and foreign firm conducting the trade and whether they are related or at arm's length, the transaction value and quantity, a ten-digit Harmonized System (HS10) code classifying the product traded, and the country of origin of the exporter.[16]

We refine the data as follows. First, we drop all transactions that are warehouse entries, so that our dataset represents imports for consumption. Second, we remove all transactions that do not include an importer identifier, an exporter identifier, an HS code, a value, a quantity or a valid transaction date. Third, we use the procedure suggested by Pierce and Schott (2012b) to create time-consistent HS codes, and correct an inconsistency in U.S. importing firms' identification codes over time by mapping firms in the LFTTD into the Longitudinal Business Database (LBD) and using the identifiers in the latter.[17] Fourth, we deflate transaction values using the quarterly GDP deflator from the FRED database maintained by the Federal Reserve Bank of Saint Louis. Finally, we collapse the refined version of the data by U.S. importer (m), foreign exporter (x), origin country (c), week the export left the foreign country (w)

[16] As noted above, import transactions are defined to be between related parties if either party owns, controls or holds voting power equivalent to 6 percent of the outstanding voting stock or shares of the other organization. We classify observations with a missing related party identifier as related. For further information on the LFTTD, see Bernard, Jensen, and Schott (2009) and Kamal, Krizan, and Monarch (2015).

[17] The inconsistency arises due to a change in single-unit firms' identification codes in 2002. We drop observations for invalid exporter identifiers, e.g., those that do not begin with a letter (it should start with the country name) or that have fewer than the requisite number of characters.

Table 1: Relationship summary statistics

	Relationship Type	
	Arm's-Length	Related-Party
Total Value Traded	$228,874$	$1,757,764$
	(11,720,829)	(79,918,870)
Overall Length (Months)	32	66
	(77)	(130)
Total Number of Shipments	4	10
	(11)	(34)
Value/Shipment (VPS)	43,257	65,379
	(601,379)	(1,091,935)
Length/Shipment (LPS)	6	10
	(15)	(22)
Number of Relationships	24,138,500	7,523,500

Notes: Table reports the mean and standard deviation of each attribute across relationships, which are defined as importer by exporter ten-digit Harmonized System category triplets observed across the 1992 to 2011 sample period. First column summarizes arm's-length relationships and second column summarizes related-party relationships (see text). Observations are restricted to relationships with more than one transaction. Value, Length, and Shipments refer to the total real value of imports observed during the relationship, the duration of the relationship in weeks, and the total number of shipments observed during the relationship. Number of observations has been rounded to the nearest 100 as per U.S. Census Bureau Disclosure Guidelines.

and ten-digit HS product (h).

We summarize the importer-exporter-product relationships observed in the data along several dimensions relevant to the model presented in the previous section. After excluding triplets with just a single shipment, we compute the total shipment value across the relationship ($Value_{mxh}$), the total length of the relationship in terms of the number of weeks between the first and last observed shipment ($Length_{mxh}$) and the total number of weeks in which a shipment occurs ($Shipments_{mxh}$) during the length of the relationship. We note that $Length_{mxh}$ is potentially subject to both left and right censoring.

The averages and standard deviations of these attributes are reported in Table 1, where the left panel contains results for arm's-length (AL) relationships and the right panel shows results for related-party (RP) relationships.[18] The table highlights that

[18]Results for AL relationships are restricted to relationships that never report an RP shipment. Results for RP relationships encompass all other relationships. We do not summarize the prices of

the average AL relationship lasts for more than two years, with shipments on average every six weeks. The large standard deviations illustrate that there is considerable variation in the length and depth of relationships.

3.2 Classifying Japanese and American Relationships

A key implication of the model of international procurement developed above is that buyers purchasing under the American system transact with a larger number of foreign sellers than buyers under the Japanese system. In this section we use this implication to classify U.S. import relationships as American or Japanese, and then investigate whether the transactions within these relationships are consistent with the model's predictions. For this exercise, we use only the arm's-length U.S. import data described in the previous section.

We classify transactions as being American or Japanese in three steps. First, we group transactions within importer by HS10 by country by mode of transportation bins in an effort to isolate likely sources of spurious variation, e.g., quality variation within HS10 products across modes of transport.[19] Then, for each bin across the entire 1992 to 2011 sample period, we compute the total number of transactions as well as the total number of distinct foreign suppliers. The ratio of these sums is the number of suppliers per shipment (SPS_{mhcz}), where z indexes mode of transportation. SPS_{mhcz} is higher when a U.S. importer uses a larger number of suppliers to obtain its imports, and has a maximum value of 1, indicating that the U.S. buyer used a different foreign exporter for every transaction within the bin. Because bins with few transactions might represent importers trying out a new product or other idiosyncrasies, we consider two classifications of bins according to whether they contain a minimum of either 5 or 15 transactions.[20] Finally, we classify an importer in an HS10 by country by mode of transport bin as American or Japanese if its SPS_{mhcz} is above or below the 90th or 10th percentiles of the supplier per shipment distribution within HS10-mode pairs across all countries for the two cutoffs, respectively.[21] Bins whose SPS_{mhcz} are above

AL vs RP relationships due to the potential influence of transfer pricing (see Bernard, Jensen, and Schott (2006)).

[19]The four main modes of transportation are vessel, rail, road, and air. We drop the small fraction of transactions that are transported by other means, e.g., hand-carried by passengers.

[20]We have also run regressions using cutoffs of $t = 10$ and $t = 20$. The results are very similar.

[21]We compute cutoffs across rather than within countries to account for the possibility that U.S. importers may choose to form Japanese relationships with suppliers from some countries but not with others. This method of computing cutoffs also allows us to obtain cross-country variation in the share

the 10th percentile but below the 90th percentile receive no classification and are not included in the first set of results presented below.

According to the model developed above, American transactions should be larger, less frequent and lower in price. Furthermore, buyer-seller relationships associated with the American system should be relatively short, due to frequent switching of suppliers. Our first approach to examining these implications is to focus on the set of importer by HS10 by country by mode of transportation bins classified as American or Japanese, and regress one of four other attributes of the bin on a dummy variable for this status: its average quantity per shipment (QPS_{mhcz}), its average number of weeks between successive shipments (WBS_{mhcz}), its average unit value per shipment ($Price_{mhcz}$), and the average length of the sales relationship between the buyer and the seller for the given product and mode of transportation ($Length_{mhcz}$). We define relationship length as the number of weeks passed between the first and the last transaction of the cell. Our second, broader approach is to use all observations in regressing bins' SPS_{mhcz} on these same attributes. Both sets of regressions include several controls. First, we include HS10-country fixed effects as well as mode of transportation fixed effects. Second, we include the total quantity transacted within the importer-HS10-country-mode cell, Qty_{mhcz}, to account for the likelihood that bins encompassing an overall larger level of imports have larger transactions or different prices due, for example, to scale effects. Including these controls allows us to compare Japanese versus American-style importers obtaining the same total quantity of the same product, from the same country and mode of transportation. We also include the weeks of the bin's first (beg_{mhcz}) and last trade (end_{mhcz}) to capture possible time and duration effects.[22]

Thus, in our first set of regressions we estimate

$$Y_{mhcz} = \beta_0 + \beta_1 d^{A5}_{mhcz} + \beta_2 \ln(Qty_{mhcz}) + \beta_3 beg_{mhcz} + \beta_4 end_{mhcz} + \lambda_{hc} + \lambda_z + \epsilon_{mhcz}, \quad (22)$$

where Y_{mhcz} is the dependent variable of interest, d^{A5}_{mhcz} is a dummy variable indicating the bin is classified as American, λ_{hc} are the product-country fixed effects, and λ_z are mode of transportation fixed effects. Standard errors are clustered at the HS10-country level. In our second set of regressions, we replace d^{A}_{mhcz} with SPS_{mhcz}.

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of relationships classified as Japanese or American.

[22]We exclude the variable beg_{mhcz} from the regression using relationship length as dependent variable, since beg_{mhcz} and end_{mhcz} jointly are highly correlated with the average relationship length. Furthermore, for importers with only one supplier for a product-country-mode, their difference is exactly the same as relationship length.

Table 2: Classification regressions at the importer level, for $t = 5$

	(1)	(2)	(3)	(4)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
d^{A5}_{mhcz}	1.221***	1.301***	−0.480***	−3.217***
	(0.003)	(0.003)	(0.005)	(0.004)
$\ln(Qty_{mhcz})$	0.756***	−0.242***	−0.355***	−0.025***
	(0.001)	(0.001)	(0.001)	(0.001)
Observations	388,000	388,000	388,000	388,000
R-Squared	0.957	0.739	0.844	0.805
Fixed Effects	hc, z	hc, z	hc, z	hc, z

	(5)	(6)	(7)	(8)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
$\ln(SPS_{mhcz})$	0.473***	0.499***	−0.185***	−1.089***
	(0.001)	(0.000)	(0.001)	(0.001)
$\ln(Qty_{mhcz})$	0.783***	−0.219***	−0.385***	−0.063***
	(0.000)	(0.000)	(0.000)	(0.000)
Observations	2,239,000	2,239,000	2,239,000	2,239,000
R-Squared	0.952	0.579	0.816	0.579
Fixed Effects	hc, z	hc, z	hc, z	hc, z

Notes: Superscripts *, **, and *** indicate statistical significance at the 10, 5, and 1 percent levels, respectively. Number of observations has been rounded to the nearest 1000 as per U.S. Census Bureau Disclosure Guidelines.

The first four columns of Table 2 report results for regressions where the key right-hand side variable is d^{A5}_{mhcz}. These regressions are restricted to bins with at least 5 transactions, using the first classification described above, but as reported in Table A.3 of Appendix B.1, results are similar for regressions restricted to bins with at least 15 transactions. As indicated in the table, we find that both the value per shipment and the number of weeks passed between shipments are more than one log point higher for bins classified as American versus Japanese. Both coefficients are also statistically significant at conventional levels. These results are not trivial: they indicate that an importer using many suppliers to import a fixed quantity receives a shipment *from any supplier* relatively less frequently and in larger lot sizes than an importer using only few exporters. As indicated in the third and fourth columns, we find statistically significantly lower transaction prices and a longer average relationship length for bins classified as American.

The bottom four columns of Table 2 report results for all bins when SPS_{mhcz} is used in place of d^{A5}_{mhcz} as the key right-hand side variable. In line with the previous columns, we find that increasing the number of suppliers per shipment by 1 percent raises the value traded per shipment by 0.47 percent, and the number of weeks between shipments by 0.50 percent. On the other hand, the average transaction price falls by 0.19 percent, and the average relationship length by 1.09 percent.

Together, the results in Table 2 indicate that classifying importers based on the number of foreign suppliers per transaction – one dimension by which American and Japanese-style procurement can be distinguished – yields results for the average order size, frequency, order price, and relationship length for the two groups that are consistent with theory. Importers purchasing the same product from many suppliers order larger lot sizes less frequently, at lower prices, and via longer relationships, while importers purchasing from few suppliers obtain inputs in smaller lot sizes, more frequently, at higher prices, and in shorter relationships.

One shortcoming of the previous analysis is that we do not control for the identity of the exporter. The model developed above predicts that suppliers under the Japanese system obtain incentive rents, which are reflected in a positive mark-up over marginal costs. Suppliers under the Japanese system should therefore charge higher prices, holding costs fixed. However, if different exporters within the same country have different costs, some suppliers might be in the Japanese system yet charge overall lower prices than suppliers under the American system due to the different cost structure. To examine this effect, we estimate equation (22) at the importer-exporter-country-product-mode level, and include exporter-product-country fixed effects. As before, the classification into American and Japanese is done for each importer-product-country-mode cell. However, the regressions now investigate whether two importers purchasing the same product from the same supplier using the same mode of transportation, but under the two different systems, differ systematically. We also introduce the additional variable $AvgLength_{mxhcz}$, which captures the average length of an importer-exporter-product-mode of transportation relationship in weeks, where the average is taken across the length of the relationship at each transaction. This variable captures the likelihood that relationships display different trading patterns when they are young than when they are old, regardless of the shipping system chosen (Heise (2015)). We exclude beg_{mxhcz} and $AvgLength_{mxhcz}$ from the regression with $Length_{mxchz}$ on the left-hand side, since relationship length is equal to the difference of beg_{mxhcz} and end_{mxhcz} and relationship length is highly correlated with its average.

Table 3: Classification regressions at the importer-exporter level, for $t = 5$

	(1)	(2)	(3)	(4)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
d^{A}_{mhcz}	0.417***	1.296***	−0.088***	−0.299***
	(0.021)	(0.040)	(0.024)	(0.032)
$\ln(Qty_{mxhcz})$	0.549***	−0.233***	−0.105***	0.218***
	(0.001)	(0.002)	(0.001)	(0.002)
Observations	183,000	183,000	183,000	183,000
R-Squared	0.981	0.790	0.969	0.719
Fixed Effects	xhc, z	xhc, z	xhc, z	xhc, z

	(5)	(6)	(7)	(8)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
SPS_{mhcz}	0.288***	0.551***	−0.077***	−0.173***
	(0.001)	(0.002)	(0.001)	(0.001)
$\ln(Qty_{mxhcz})$	0.659***	−0.176***	−0.151***	0.155***
	(0.000)	(0.001)	(0.001)	(0.001)
Observations	1,686,000	1,686,000	1,686,000	1,686,000
R-Squared	0.980	0.688	0.957	0.617
Fixed Effects	xhc, z	xhc, z	xhc, z	xhc, z

Notes: Superscripts *, **, and *** indicate statistical significance at the 10, 5, and 1 percent levels, respectively. Number of observations has been rounded to the nearest 1000 as per U.S. Census Bureau Disclosure Guidelines.

The general specification is:

$$\begin{aligned} Y_{mxhcz} &= \beta_0 + \beta_1 d^{A}_{mhcz} + \beta_2 \ln(Qty_{mxhcz}) + \beta_3 beg_{mxhcz} + \beta_4 end_{mxhcz} \\ &\quad + \beta_5 \ln(AvgLength_{mxhcz}) + \lambda_{xhc} + \lambda_z + \epsilon_{mxhcz}. \end{aligned} \tag{23}$$

For this regression, standard errors are clustered at the importer-HS10-country level.

Results for $t = 5$, presented in Table 3, are consistent with those in Table 2: an importer whose number of suppliers is in the 90th percentile of the SPS distribution purchases on average 41 percent more quantity per shipment and receives shipments spaced on average more than twice as far apart versus an importer who purchases from the same supplier but is in the 10th percentile of the SPS distribution. Moreover, the supplier charges a price that is on average 8.8 percent lower to an importer who we classify as procuring under the American system versus an importer classified as using the Japanese system, and the relationship is 30% shorter. Results using SPS_{mhcz}

directly are similar here, too, as are those for bins with at least 15 transactions (see Table A.4 in Appendix B.1).

Given these results within a given product category, we next proceed to study whether we can detect systematically different shipment patterns in the cross-section of products. Rauch (1999) classifies products into whether they are traded on organized exchanges, have reference prices, or neither. The latter products are referred to as differentiated. If differentiated products are more costly to inspect, we should observe that such products exhibit shipment patterns that are more consistent with the Japanese system, while reference-priced products display more American patterns and products traded on organized exchanges are again more American. Since the classification into American and Japanese assigns relationships within each product based on percentiles, we cannot simply compare the share of Japanese relationships within products. Instead, we collapse our dataset into importer-product-country-mode of transportation cells as before, and compare the value per shipment (VPS_{mhcz}), the average number of weeks between successive shipments (WBS_{mhcz}), and the average length of the sales relationship ($Length_{mhcz}$) across products. We do not include the average price or quantity per shipment in these regressions, since prices and quantities vary across products for many reasons. Our regression specification is

$$Y_{mhcz} = \beta_0 + \beta_1 d_h^D + \beta_2 d_h^{Ref} + \beta_3 \ln(Value_{mhcz}) + \beta_4 beg_{mhcz} + \beta_5 end_{mhcz} + \lambda_{mc} + \lambda_z + \epsilon_{mhcz}, \tag{24}$$

where d_h^D is a product-level dummy for differentiated products, and d_h^{Ref} is a dummy for products that have reference prices. We control for the total (deflated) value traded of an importer-product-country-mode cell across the entire dataset, and include the weeks of the bin's first (beg_{mhcz}) and last trade (end_{mhcz}) as well as importer-country and mode of transportation fixed effects. The identification of this regression comes from importers purchasing from the same country under the same mode of transportation products that are classified differently by Rauch (1999). We include only cells with at least $t = 5$ transactions. Standard errors are clustered at the product-level.

The first three columns of Table 4 present the regression coefficients on the two Rauch dummies, using the liberal specification developed by Rauch (1999). We find that differentiated products exhibit a smaller value per shipment (column 1) and more frequent shipments (column 2) than products that have reference prices, in line with more Japanese-style shipment patterns. Furthermore, the reference-priced products are shipped in smaller lot sizes and more frequently than products that are traded

Table 4: Rauch Regressions for $t = 5$ (Liberal Classification)

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	ln(VPS)	ln(WBS)	ln(Length)	ln(VPS)	ln(WBS)	ln(Length)
d_h^D	-0.198^{***}	-0.195^{***}	0.135^{***}	-0.161^{***}	-0.164^{***}	0.185^{***}
	(0.003)	(0.004)	(0.006)	(0.003)	(0.004)	(0.005)
d_h^{Ref}	-0.120^{***}	-0.118^{***}	0.205^{***}	-0.137^{***}	-0.132^{***}	0.183^{***}
	(0.003)	(0.004)	(0.006)	(0.003)	(0.004)	(0.006)
$\ln(NSupp_{mhcz})$				-0.456^{***}	-0.374^{***}	-0.608^{***}
				(0.001)	(0.001)	(0.002)
Observations	$1,556,000$	$1,556,000$	$1,556,000$	$1,556,000$	$1,556,000$	$1,556,000$
R-Squared	0.874	0.700	0.532	0.904	0.727	0.585
Fixed Effects	mc, z	mc, z	mc, z	mc, z	mc, z	mc, z

Notes: Superscripts *, **, and *** indicate statistical significance at the 10, 5, and 1 percent levels, respectively. Number of observations has been rounded to the nearest 1000 as per U.S. Census Bureau Disclosure Guidelines.

on organized exchanges. The overall relationship length for differentiated products is below the one for products with reference prices, but both product categories have longer relationships than products traded on organized exchanges (column 3). Thus, shipping patterns across products are well in line with our model predictions.

One issue with the previous specification is that importers purchasing differentiated products are likely to have relatively more suppliers per good due to the existence of more varieties. Having more suppliers could lead to smaller shipments from each supplier and to shorter relationships with each. To control for this effect, the last three columns of 4 include the total number of suppliers for the importer-product-country-mode bin ($NSupp_{mhcz}$) as additional regressor. We find that controlling for the number of suppliers, differentiated products have a slightly longer average relationship length than products with reference prices, and exhibit smaller, more frequent shipments, as expected. Table A.2 in Appendix B.1 shows our results under the conservative Rauch classification. Our findings are further strengthened under that specification. Overall, the results provide significant support for our theory and show that it can also explain shipment patterns in the cross-section.

3.3 The Effect of PNTR on the Choice of Procurement System

The model presented in Section 2 suggests that the share of American and Japanese procurement relationships in the economy can vary with trade policy. In particular, an increase in the probability of peaceful trade can induce buyer and seller to switch from the American to the Japanese system. Such a switch may lower procurement costs and increase consumer welfare.

We study this implication using a plausibly exogenous change in U.S.trade policy, the U.S. granting of PNTR to China in October 2000, which substantially reduced the possibility of a trade-war-like hike in U.S. import tariffs on Chinese goods. U.S. imports from non-market economies such as China are generally subject to non-NTR tariff rates originally set under the Smoot-Hawley Tariff Act of 1930. These rates typically are substantially larger than the NTR rates the U.S. offers fellow members of the World Trade Organization (WTO) – 32 percentage points higher on average, as discussed below. The U.S. Trade Act of 1974 allows the President to grant NTR tariff rates to non-market economies on an annually renewable basis subject to Congressional approval, and U.S. Presidents began granting such a waiver to China in 1980. While these waivers kept the actual tariff rates applied to Chinese goods low, the need for annual approval by Congress created uncertainty about whether the low tariffs would continue, particularly during the 1990s. In our model, this implies that the probability that relationships with U.S. firms can continue depends on the uncertainty that NTR tariffs are renewed and the amount that tariffs would increase if China's NTR status was withdrawn. A low probability of renewal associated with a high pre-NTR tariff results in a lower probability that U.S.-China relationships are long lasting.

Summary Statistics in the Data

We begin to assess the effect of PNTR on the structure of supply chains by plotting the share of Japanese-style relationships in U.S.-China trade over time using an approach similar to the one described in the previous section. First, we divide the sample period into four time windows: 1992 to 1996, 1997 to 2001, 2002 to 2006, and 2007 to 2011. Then, for each of these four windows, we compute the number of suppliers per shipment, SPS_{mhcz}, for each importer, HS10, country and mode of transportation bin with at least 5 transactions. We then use the 10th percentile of the SPS-distribution across

Table 5: Share of Japanese-Style Relationships by Country, 1992-2001

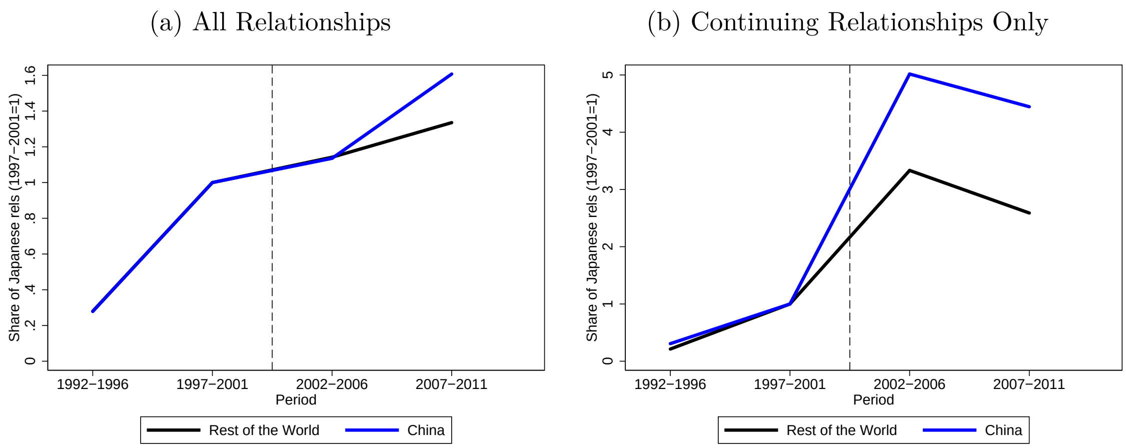
Country	Share of Japanese	Country	Share of Japanese
Mexico	14.27%	United Kingdom	8.33%
Japan	11.27%	South Korea	8.18%
Germany	9.63%	Brazil	7.50%
Rest of the World	9.50%	France	6.29%
Canada	9.16%	China	5.32%
Taiwan	8.92%		

importers and countries for each HS10-mode pair for the second time period (1997 to 2001, i.e., the period just before the change in trade policy) to classify bins as Japanese in all time periods. Note that while 10 percent of bins are classified as Japanese during the second time period by construction, the share of bins classified as Japanese in the other windows can vary. Finally, we compute the value-weighted average share of bins that are Japanese across HS10 codes and modes of transportation for each window, both for U.S. imports from China and for U.S. imports with the rest of the world.

Table 5 provides some details on the share of Japanese-style relationships for U.S. trade with its 10 largest trading partners and with the rest of the world just prior to the policy change. For this table, we compute a value-weighted average over the share of relationships classified as Japanese in the first two time windows (1992-1996 and 1997-2001), for each country. We find that Japanese-style relationships account for approximately 5.3 percent of all U.S.-China imports in the pre-PNTR period. This share is significantly smaller than for any other country among the largest trading partners. We expect trade with Japan to exhibit a high share of Japanese-style relationships, since firms such as Toyota are well known to have close relationships with suppliers and customers, and that is what we find in the data. Mexico has the highest share of Japanese-style relationships, possibly due to the close supply chain integration with the U.S.

Figure 7a shows the evolution of the share of Japanese-style relationships in total trade relative to the 1997 to 2001 period, which we normalize to one. As indicated in the figure, the share of Japanese relationships increases over time. In the 2002 to 2006 period, which immediately follows PNTR, the share of Japanese relationships grows by about 20 percent compared to the 1997 to 2001 period, both for U.S.-China trade and for trade with the rest of the world. While our framework predicts a relatively faster increase in the share of Japanese relationships with China due to PNTR, the

Figure 7: Share of Japanese-Style Relationships in U.S. Trade



lack of an immediate effect might be due to a large number of U.S. importers exploring importing from China and forming new relationships after 2001.[23] We find that the share of Japanese relationships grew significantly more rapidly for trade with China than for trade with the rest of the world during the 2007 to 2011 period, to about 61 percent and 34 percent above the baseline level in 1997 to 2001, respectively.

To remove the effect of a possible surge in relationship formation following 2001, we compute the same figure for only importer-exporter-product triplets that exist both before PNTR (trading at least once before 2001) and after PNTR (trading at least once from 2002 onwards). Focusing on this sub-sample allows us to examine how PNTR affects continuing relationships. Figure 7b highlights that there is a substantial increase in the share of continuing Japanese-style relationships with China in the period immediately following PNTR, and this increase is much stronger than the increase for the rest of the world. The share of continuing Japanese-style relationships with China increased by a factor of 5 in the 2002-2006 period compared to the period 1997-2001, indicating a substantial switch towards Japanese-style procurement.

Difference-in-Difference Analysis

To assess the effects of PNTR on procurement patterns for imports from China more carefully, we perform a differences-in-differences regression. We define the NTR gap

[23] In fact, we show below that after PNTR a large number of new relationships were formed with Chinese suppliers and these were often short-lived.

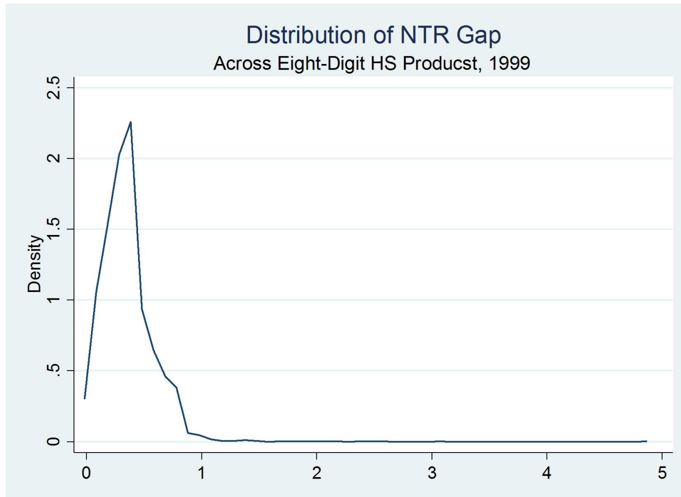


Figure 8: Distribution of the NTR Gap Across Import Products

for eight-digit HS import product h as the difference between non-NTR and NTR tariff rates,

$$NTR\,Gap_h = Non\ NTR\ Rate_h - NTR\ Rate_h, \tag{25}$$

using *ad valorem* equivalent tariff rates provided by Feenstra, Romalis, and Schott (2002) for 1999, the year before passage of PNTR in the United States.[24] As indicated in Figure 8, these gaps vary widely across products, and have a mean and standard deviation of 0.32 and 0.23. Our identification strategy exploits this variation in the NTR gap to determine whether U.S.-China procurement patterns change relative to procurement patterns with exporters from other source countries (first difference) after the change in U.S. policy is implemented (second difference) in industries with higher NTR gaps (third difference). The last difference captures the fact that industries with larger NTR gaps experience a larger increase in the relationship continuation probability than industries with smaller gaps. We expect the largest shifts toward Japanese-style procurement after PNTR to occur in U.S. imports of high-gap products from China.

Our first, preferred specification compares shipments within importer-exporter-product triplets across two symmetric time intervals around the change in U.S. trade policy, $p \in \{Pre, Post\}$,

[24] While U.S. tariffs are set at the level of eight-digit HS products, we observe trade at the ten-digit HS level. In our empirical work, we therefore match each ten-digit HS product with the tariff associated with its first eight digits.

$$
\begin{aligned}
ln(\overline{Y}_{mxhcp}) =& \beta_0 + \beta_1 1\{p = Post\} * 1\{c = China\} * NTR\,Gap_p + \gamma\chi_{mxchp} \\
&+ \beta_2 ln(Total\,Value_{mxhcp}) + \lambda_{mxh} + \lambda_c + \lambda_p + \epsilon_{mxhcp}
\end{aligned}
\tag{26}
$$

where subscripts m, x, h and p index U.S. importers, exporters from country c, ten-digit HS products and time period. The regression sample consists of all shipments by "always-arm's-length" parties, i.e., parties that engage solely in arm's length transactions over the entire 1992 to 2011 sample period, so long as there is at least one shipment in each period. Periods are one of two distinct five-year windows around 2001, either 1995 to 2000 (pre period) or 2002 to 2007 (post period). Note that the latter window ends before the Great Recession, and also before we observe the largest increase in the share of Japanese-style relationships with China in the simple plot in Figure 7a.

$\overline{Y}_{mxhcp}$ represents one of several attributes of shipment patterns within an $mxhcp$ bin deemed relevant by the model developed in Section 2: WBS_{mxhcp} is the average number of weeks between shipments, VPS_{mxhcp} is the average value per shipment, QPS_{mxhcp} is the average quantity per shipment, $Price_{mxhcp}$ is the average unit value per shipment, and $Length_{mxhcp}$ is the average length in weeks of the importer-exporter-product relationships appearing within the $mxhcp$ bin.[25] The matrix χ_{mxhcp} represents the full set of interactions of the NTR gap, the post dummy variable ($1\{p = Post\}$) and the China dummy variable ($1\{c = China\}$) required to identify β_1. $TotalValue_{mxhcp}$ is the total value of all shipments occurring within the $mxhcp$ bin; its inclusion accounts for the varying scale of imports across bins. Relationship (mxh), country and period fixed effects are represented by δ_h, δ_c and δ_p. The difference-in-differences coefficient of interest, β_1, measures the log difference in activity for shipments from China versus other countries after the change in U.S. policy versus before for products with higher versus lower NTR gaps. From the model presented in Section 2, we expect $\beta_1 < 0$ for VPS_{mxhcp}, QPS_{mxhcp} and WBS_{mxhcp}, and $\beta_1 > 0$ for $Price_{mxhcp}$ and $Length_{mxhcp}$ if PNTR induced a switch from the American to the Japanese system.[26]

[25]The length of each relationship is defined as the number of weeks between the first observed transaction during the period and the last observed transaction during the period.

[26]There are several motivations for why PNTR might *not* have induced procurement patterns to become more Japanese. For example, if PNTR caused only a small increase in Chinese exporters' assessment trade peace, switching from American to Japanese procurement would be minimal. As a result, import patterns for American-style importers would remain unchanged while those for Japanese-style

The second specification ignores exporter identity and analyzes shipments within importer-products across periods,

$$\begin{aligned} ln(\overline{Y}_{mhcp}) =& \beta_0 + \beta_1 1\{p = Post\} * 1\{c = China\} * NTR\,Gap_p + \gamma\chi_{mhcp} \\ &+ \beta_2 ln(Total\,Value_{mhcp}) + \delta_{mh} + \delta_c + \delta_p + \epsilon_{mhcp} \end{aligned} \tag{27}$$

Here, too, the regression sample includes all shipments by "always-arm's-length" parties so long as there is at least one shipment for each *mhcp* bin. After the procurement attributes are computed, the *mxhcp* data are collapsed to the *mhcp* level so that there is one observation – the average – in the regression for each *mhcp* bin.

Our final specification ignores both importer and exporter identity and analyzes shipments within products across periods,

$$\begin{aligned} ln(\overline{Y}_{hcp}) =& \beta_0 + \beta_1 1\{p = Post\} * 1\{c = China\} * NTR\,Gap_h + \gamma\chi_{hcp} \\ &+ \beta_2 ln(Total\,Value_{hcp}) + \delta_h + \delta_c + \delta_p + \epsilon_{hcp} \end{aligned} \tag{28}$$

As above, we require at least one shipment within each *hcp* bin, and the data are collapsed to the *hcp* level after the procurement attributes are computed.

Results for the first, second and third specifications are reported in the corresponding three columns of Table 6, where each row reports the estimated DID term coefficient and standard error for a different relationship attribute. Starting with the preferred, within-*mxh* results reported in column 1, we find that all estimates of β_1 are consistent with a switch towards Japanese procurement: point estimates for value per shipment, quantity per shipment and weeks between shipments are all negative, though statistically significant only for the first two, while they are positive and statistically significant for shipment price and overall length. In terms of economic significance, these results imply that a one standard deviation increase in the NTR gap (0.23) is associated with relative declines in shipment value and shipment quantity of 1.6 and 3.0 percent after the change in U.S. policy. Shipment price and relationship length, by contrast, rise by 0.9 and 2.3 percent, respectively.

importers would become slightly larger at a slightly higher price. Or, the most salient impact of PNTR and China joining the WTO might have been to give U.S. importers greater confidence in the enforceability of contracts. In that case, a shift from Japanese to American might be expected.

Table 6: PNTR and procurement

Dependent Variable	Within Importer-Exporter-Product	Within Importer-Product	Within Product
ln(Value per Shipment)	-0.07^{***}	-0.05^{***}	-0.17^{***}
	0.01	0.01	0.05
ln(Quantity per Shipment)	-0.13^{***}	-0.04^{**}	-0.04
	0.02	0.02	0.10
ln(Price per Shipment)	0.04^{**}	-0.04^{**}	-0.11
	0.02	0.02	0.09
ln(Weeks between Shipments)	-0.04	-0.06^{***}	-0.36^{***}
	0.03	0.02	0.07
ln(Overall Relationship Length)	0.10^{***}	0.00	-0.34^{***}
	0.04	0.03	0.08
Observations	$752,600$	$1,011,700$	$324,300$
Sample	*mxhcp*	*mhcp*	*mcp*
Fixed Effects	mxh, c, p	mh, c, p	h, c, p

Notes: Table summarizes the results of generalized differences-in-differences regressions of relationship attributes on a DID coefficient representing the interaction of the NTR gap and dummy variables representing the post-PNTR period and trade with China (see text). Each cell in the table represents the result of a different regression. Data are collapsed to the importer-exporter-product-country-period (*mxhcp*) level in column 1, the importer-product-country-period (*mhcp*) level in column 2, and the product-country-period (*hcp*) level in column 3. Sample is restricted to bins with at least two observations for the pre- and post period. Dependent variables are computed with respect to the noted sample bins. Results for fixed effects and other covariates needed to identify the DID coefficient of interest are suppressed. Superscripts *, **, and *** indicate statistical significance at the 10, 5, and 1 percent levels, respectively. Number of observations has been rounded to the nearest 100 as per U.S. Census Bureau Disclosure Guidelines.

Comparison of the within-relationship results in column 1 with the within-product results in column 3 provides further intuition for our theoretical framework, and illuminates our findings in the initial analysis of the share of Japanese-style relationships. For example, the relatively large (in absolute terms) DID point estimates for VPS_{hcp}, WBS_{hcp} and $Length_{hcp}$ reflects the fact that the change in U.S. policy gave rise to many new relationships. Since many of these relationships involved firms that had not imported from China before (see Pierce and Schott 2016), it is unsurprising that they were short-lived and perhaps encompass smaller, trial-size shipments.

Table 7: Simulation Parameters

	Baseline
Interest Rate (r)	0.02
Elasticity of Substitution (σ)	4
High Quality ($\bar{\theta}$)	10
Trade war with self ($\Psi_{i,i}$)	1
Trade war U.S.-RoW ($\Psi_{US,RoW}$)	0.98

4 Quantitative Analysis

In this section we quantify the welfare gains due to changes in procurement driven by an exogenous increase in the probability of trade peace. These welfare effects have two sources. First, the lower costs of Japanese-style procurement from China lead to a switch towards this procurement system for products already sourced from China under the American system. Second, new trade relationships are formed for products where China is now the lowest cost origin country.

We first discuss how we estimate the model parameters. We then simulate the model to assess the implications of PNTR. Since the model can only be solved using numerical methods, we consider a setup with $N = 3$ countries and $J = 100$ products, where $n = 1$ represents the U.S., $n = 2$ is China, and $n = 3$ is the Rest of the World.

4.1 Estimation and Identification

Estimation Strategy

We set a number of the model's parameters based on existing literature. First, we consider an annual time horizon, and set $r = 0.02$. We choose an elasticity of substitution $\sigma = 4$ as in Nakamura and Steinsson (2008), which delivers a mark-up of buyers for their final goods close to estimates by Berry, Levinsohn, and Pakes (1995). For the variable costs, we set $\bar{\theta} = 10$. We assume that the per-period probabilities of trade peace, $\Psi_{ni} = e^{-\rho_{ni}}$ are symmetric, and set own-country probabilities to one. Furthermore, we assume that a trade war between the U.S. and the Rest of the World is unlikely, and set the annual probability of trade peace to $\Psi_{US,RoW} = 0.98$. These parameters values are presented in Table 7.

Three sets of parameters remain to be estimated: the shipment cost parameters f^j and d^j_{ni}, the inspection costs m^j, and the remaining per-period probabilities of trade peace, Ψ_{ni}. We estimate these parameters via a simulated method of moments procedure using moments observed in the LFTTD data and in external data. While the parameters are jointly estimated, we proceed to describe the empirical moments targeted and the underlying identification assumptions for each parameter in turn.

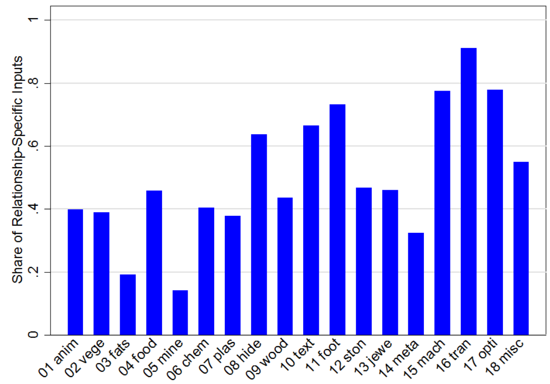
The first set of parameters to be estimated are the remaining probabilities of trade peace. To set the pre-PNTR probability of trade peace between the U.S. and China, $\Psi^{pre}_{US,China}$, we target the observed share of Japanese-style relationships in U.S.-China trade, which based on Table 5 is about 5.3%. Since the share of Japanese relationships in US-Chinese trade is strictly increasing in $\Psi_{US,China}$, this parameter is well identified for a given level of inspection costs, which we set below. We set the post-PNTR probability of trade peace, $\Psi^{post}_{US,China}$, to generate the increase in the share of Japanese-style trade between the U.S. and China that we observe in the data post PNTR. To account for the fact that the share of Japanese relationships is generally increasing across all countries, we target only the differential increase in Japanese-style relationships with China relative to the Rest of the World. Based on our previous results, we obtain an increase in the share of Japanese-style relationships due to PNTR of about 27%. We set the probability of trade peace between China and the Rest of the World to the same level as $\Psi^{pre}_{US,China}$.

We estimate parameter values for m^j, f^j, and d^j_{ni} for 18 broad product categories, which are based on the first two digits of the HS10 code.[27] Thus, we estimate moments for 18 tuples of the form $(m^j, f^j, \left\{d^j_{ni}\right\}_{n,i})$ based on our transaction-level data. In the simulation, we will then take J draws from these product bins with probability weights in proportion to their value share in U.S. imports..

We estimate the 18 inspection cost parameters m^j using a two-step procedure. First, we estimate the average inspection cost $\bar{m}$ using the average share of Japanese relationships observed in the LFTTD across all countries in the pre-PNTR period. This share is about 8.6% from Table 5. The parameter is well-identified since the total share of Japanese relationships was not used in the estimation of the Ψ, which used the share of Japanese relationships in U.S.-China trade only. In the second step, we obtain a distribution for m^j using the measure of contract intensity by Nunn (2007). His paper estimates for each 6-digit BEA industry code the share of intermediate inputs

[27]These categories are: 1. Animal products (HS2=01-05), 2. Vegetables (06-14), 3. Fats (15), 4. Food (16-24), 5. Minerals (25-27), 6. Chemicals (28-38), 7. Plastics (39-40), 8. Hide (41-43), 9. Wood (44-49), 10. Textiles (50-63), 11. Footwear (64-67), 12. Stones (68-70), 13. Jewelry (71), 14. Metals (72-83), 15. Machinery (84-85), 16. Transportation (86-89), 17. Optics (90-92), 18. Miscellaneous (94-96).

Figure 9: Relationship Specificity Distribution



that is relationship-specific, where specificity is defined as the share of inputs that is neither sold on an organized exchange nor reference priced, based on the classification by Rauch (1999). Our assumption is that products that exhibit a higher degree of relationship specificity are more likely to be complex, which makes them more costly to inspect. We use the liberal classification measure provided, and map BEA codes to NAICS codes and from there to HS10 codes using the concordance by Pierce and Schott (2012a). We then aggregate the shares of relationship-specific inputs to the 18 HS2 product categories, taking a value-weighted average using the import value of each industry in 2002 from the U.S. Census.

Figure 9 shows the estimated share of relationship-specific inputs by product category. The value-weighted average share across all categories of 0.616. To obtain the distribution of m^j, we calculate the ratio between a category's relationship specificity and the mean of 0.616, and apply these ratios to the mean inspection cost $\bar{m}$.

We pin down the shipment cost parameters f^j and d^j_{ni} for the 18 product categories using the shipping frequencies observed in the data. Previous work seeking to estimate how distance affects shipment costs such as Limao and Venables (2001) and Hummels (2007) has used cost data such as shipping company quotes or air fares to estimate the elasticity of transport costs with respect to distance and other covariates. Here, we take a different approach and use the frequency of shipments to estimate an elasticity of shipment costs with respect to distance. We exploit the fact that in our model, higher values of f^j and d^j_{ni} are associated with less frequent, larger shipments (Proposition 4). Thus, we can use variation in the number of weeks between shipments for importer-exporter pairs transacting the same total value of the same HS10 good via the same mode of transportation from different countries to determine a good-specific cost component that is independent of distance, f^j, and a distance-specific part d^j_{ni}. To the extent that these distance elasticities are positive, and hence a larger distance implies less frequent shipments, this estimation provides further support of one of the implications of our model.

We proceed in three steps. First, for each of the 18 product categories, we run a regression of the form

$$\begin{aligned} \ln(WBS_{mhcz}) = \beta_0 + \beta_1 \cdot \ln(Dist_c) + \beta_2 \cdot \ln(Quantity_{mhcz}) + \beta_3 \cdot \ln(SPS_{mhcz}) \\ + \lambda_h + \lambda_z + \epsilon_{mhct}, \end{aligned} \tag{29}$$

where WBS_{mhcz} is the average number of weeks between shipments for a given importer m purchasing HS10 product h from country c using mode of transportation z, $Dist_c$ is the great circle distance between the most important city / agglomeration in the exporting country and in the U.S. provided by the CEPII GeoDist database, $Quantity_{mhcz}$ is the total quantity purchased by the importer-product-country-mode cell, λ_h are HS10 fixed effects, and λ_z are mode of transportation fixed effects. Running the regression for each product category separately seeks to account for variation in shipment patterns that is due to product-specific factors.[28] Based on our result that shipping patterns are sensitive to the procurement system used, we also include controls for suppliers per shipment (SPS) to compare shipments within a given procurement system. The regression coefficients for β_1 and β_3 and their standard errors are presented in the first two columns of Table 8. Standard errors are clustered at the HS10-country level. We find a significant and positive effect of distance on the number of weeks between shipments for all product categories, with the exception of animal products and minerals. Overall, an increase in the distance from the U.S. by 1% raises the number of weeks between shipments by 0.09%. Thus, distance has a positive effect on shipment frequency as predicted. As expected, we also find that an increase in suppliers per shipment increases the time gap between shipments.

As the second step, we construct the predicted value of $\ln(WBS_{mhcz})$ for each transaction under the assumption that $\ln(Dist_c) = 0$ and SPS is at the 10th percentile for

[28] Limao and Venables (2001) and Hummels (2007) point out other variables affecting shipment costs, such as the weight of the item shipped. The split into different product categories seeks to capture variation across products in this dimension.

Table 8: Frequency regression results

HS2 and product category	Dist (β_1)	SPS (β_3)	HS2 and product category	Dist (β_1)	SPS (β_3)
All	0.088***	0.499***			
	(0.001)	(0.001)			
01-05	−0.017***	0.454***	50-63	0.133***	0.418***
Animal prod.	(0.006)	(0.005)	Textiles	(0.002)	(0.001)
06-14	0.012*	0.406***	64-67	0.221***	0.438***
Vegetables	(0.007)	(0.004)	Footwear	(0.007)	(0.004)
15	0.056**	0.489***	68-70	0.111***	0.460***
Fats	(0.025)	(0.013)	Stones	(0.008)	(0.004)
16-24	0.059***	0.398***	71	0.062***	0.488***
Food	(0.006)	(0.003)	Jewelry	(0.018)	(0.014)
25-27	−0.026***	0.579***	72-83	0.081***	0.520***
Minerals	(0.011)	(0.009)	Metals	(0.003)	(0.002)
28-38	0.008	0.484***	84-85	0.111***	0.550***
Chemicals	(0.005)	(0.003)	Machinery	(0.003)	(0.002)
39-40	0.105***	0.515***	86-89	0.097***	0.525***
Plastic	(0.005)	(0.003)	Transportation	(0.006)	(0.004)
41-43	0.273***	0.466***	90-92	0.228***	0.509***
Hide	(0.007)	(0.004)	Optics	(0.006)	(0.003)
44-49	0.067***	0.441***	93-99	0.181***	0.498***
Wood	(0.005)	(0.003)	Miscellaneous	(0.006)	(0.003)

the product category.[29] This estimate, which we call $\ln(W\hat{B}S^{0}_{mhcz,10})$, is the counterfactual shipment frequency under the Japanese system if the distance for a given transaction had been zero. We similarly construct $\ln(W\hat{B}S^{0}_{mhcz,90})$ for the implied shipment frequency at the 90th percentile (corresponding to the American system). By taking a value-weighted average of these two variables across all transactions within each product category, we obtain product category-specific counterfactual shipment frequencies at distance zero under the Japanese system and under the American system. These shipment frequencies are shown in Table 9. We set the f^j to match these 36 moments as closely as possible. We find that shipments are relatively infrequent in particular for minerals, chemicals, and animal products, and hence the fixed cost of preparing a production run seems to be relatively high for these products. On the

[29] We assume that $\beta_1 = 0$ for those product categories where the distance elasticities are estimated to be negative.

Table 9: Frequency moments

HS2 code	Product category	$\hat{WBS}_A^0$	$\hat{WBS}_J^0$
	All	8.971	2.232
01-05	Animal products	11.747	3.075
06-14	Vegetables	9.874	3.162
15	Fats	10.799	2.841
16-24	Food	7.223	2.326
25-27	Minerals	34.983	5.563
28-38	Chemicals	22.835	5.899
39-40	Plastic	7.622	1.628
41-43	Hide	1.258	0.379
44-49	Wood	8.247	2.440
50-63	Textiles	4.091	1.418
64-67	Footwear	1.309	0.413
68-70	Stones	6.102	1.778
71	Jewelry	6.395	1.534
72-83	Metals	8.770	1.967
84-85	Machinery	9.121	1.840
86-89	Transportation	6.858	1.335
90-92	Optics	2.634	0.641
93-99	Miscellaneous	2.600	0.699

other hand, shipping occurs frequently for footwear, hide products, and optics, where the fixed costs are consequently estimated to be low.

The third step consists of the estimation of the distance costs d^j_{ni}. We compute the distance between the U.S. and the Rest of the World as a value-weighted average distance between the largest agglomeration in the U.S. and in each exporting country, using the value of imports in 2002 from the U.S. Census as weights. We similarly estimate the distance between China and the Rest of the World using trade flows from the UN COMTRADE database as weights. For the distance of the Rest of the World with itself, we calculate the value-weighted distance between all country pairs. These distances, together with the results from regression (29), imply estimated shipment frequencies for each of the 18 product categories at each distance. We generate the frequencies again for both the American and the Japanese system, which are set as the shipment frequencies at the 10th and the 90th percentile of the SPS distribution. We choose the d^j_{ni} to match these implied shipment frequencies under each system.

Our procedure targets in total 2 moments for probabilities of trade peace, one moment for inspection costs, 36 moments for counterfactual shipping probabilities at distance zero, and $2 \times 4 \times 18 = 144$ moments for distances, which sums to a total of 183 moments. Let the true values of the parameters in the data be Θ, and denote the estimated parameters by $\hat{\Theta}$. We denote the vector of data moments and model moments by $G(\Theta)$ and $G(\hat{\Theta})$, respectively. We estimate a total of 93 parameters by choosing the parameter values that minimize

$$J = \min_{\hat{\Theta}} E\left[(G(\Theta) - G(\hat{\Theta}))'(G(\Theta) - G(\hat{\Theta}))\right]. \tag{30}$$

Estimation Results

Work in progress. We can show our distance cost estimates and how they compare to earlier work, etc.

4.2 Effects of a Change in the Probability of Trade Peace

[The results here are illustrative only] Given the estimated model, we are now in a position to simulate the effect of PNTR on U.S. trade flows by increasing the probability of trade peace from $\Psi^{pre}_{US,China}$ to $\Psi^{post}_{US,China}$. Table 10 presents the results of these simulations. The first three columns present statistics before PNTR, for the U.S., China, and the Rest of the World. The last three columns of the table refer to the same countries post-PNTR.

The first two rows of Table 10 illustrate the trade diversion effects of PNTR. We compute the fraction of the value imported into each region from China,

$$V_{nc} = \frac{\int p^j_{nc} q^j_{nc}}{\int p^j_{ni} q^j_{ni}},$$

where $i = c$ if the origin country is China. Prior to PNTR, the U.S. imports about half of its value from the rest of the world, and nothing from China. PNTR lowers the costs of imports from China under the Japanese system, leading some products to be switched towards Chinese suppliers. This switch raises the share of value imported from China to nearly one fifth, at the expense of imports from the rest of the world, which fall by 15 percentage points, and at the expense of domestic U.S. suppliers. The consequences of the switch of domestic production towards imports from China have been documented in Pierce and Schott (2016).

The shift towards the Japanese system is illustrated in the third row of Table 10. The value share of products sourced by U.S. importers under the Japanese system is about 46% pre-PNTR, but rises to 51% afterwards, across all countries. In line with this shift towards Japanese-style procurement, the average number of shipments, for the average product, rises in the U.S. from 0.443 shipments per period to 0.489 shipments per period (row 4). Rows 5-6 show that this increase in shipment frequency is almost solely due to products switching from American- to Japanese-style procurement. The small increase in for non-switchers is due to the increase in total quantity traded, q_{ni}^{j}, which makes more frequent shipments less costly. Since the probability of trade peace with China does not change for the Rest of the World, these countries do not experience a shift in their procurement patterns.

The effects of less costly procurement from China on the U.S. consumer price index is presented in row 7 of Table 10. All prices are expressed relative to the U.S. pre-PNTR price index. We find that consumer prices in the U.S. fall by 0.2% as a result of PNTR. There is also a small decrease in the Chinese price index, as Chinese importers are now also able to procure more easily under the Japanese system from the U.S. The flipside of this movement is an increase in aggregate consumption (row 8). Aggregate U.S. consumption, and hence consumer welfare, rises by 0.2% as a direct consequence of PNTR. This result highlights that non-traditional trade policies on their own can generate welfare increases, in the case of PNTR as a result of differences in the likelihood of maintaining relationships, even without the presence of fundamental differences in productivity.

The last row of Table 10 shows the effect of PNTR on real wages. Since labor supply is one in each country, the real wage is exactly equal to the total quantity consumed. Note that the wage level in China is slightly below the wage level in the U.S. The relatively high probability of a trade war with China deters other countries from importing Chinese goods, which necessitates a drop in Chinese real wages to stimulate exports in order to clear the labor market. However, while Chinese exporters are hurt by the higher likelihood of trade wars with the U.S., they also benefit from the fact that by symmetry it is also more difficult for U.S. exporters to export to China. Since there are no cross-country productivity differences in the model, wages in the more closed Chinese economy remain relatively similar to the wages in the U.S. If productivity differences were introduced into the model by choosing a lower value of Υ_i for China, it would become easier for foreign firms to break into the Chinese

Table 10: Simulated Effects of PNTR

	Before PNTR			**After PNTR**		
	United States	China	Rest of World	United States	China	Rest of World
Value from U.S. (%)	59.7%	12.4%	19.9%	51.7%	28.1%	19.9%
- of which, "Japanese"	79.2%	0.0%	0.0%	76.1%	55.8%	0.0%
Value imported from China (%)	23.8%	59.3%	11.7%	31.9%	51.4%	11.7%
- of which, "Japanese"	0.0%	86.3%	0.0%	74.0%	84.2%	0.0%
Value imported from ROW (%)	16.5%	28.3%	68.4%	16.4%	20.6%	68.4%
- of which, "Japanese"	0.0%	0.0%	2.5%	0.0%	0.0%	2.5%
Number of shipments per period	0.513	0.538	0.493	0.599	0.579	0.493
- Products switching to J system	0.236	0.235	–	0.752	0.717	–
- Products that do not switch	0.566	0.564	0.493	0.569	0.567	0.493
Aggregate price index (U.S.=1)	1.000	1.000	0.989	0.998	1.000	0.989
Aggregate quantity (U.S.=1)	1.000	1.000	1.011	1.002	1.000	1.011
Real wage level (U.S.=1)	1.000	1.000	0.989	1.002	1.000	1.011

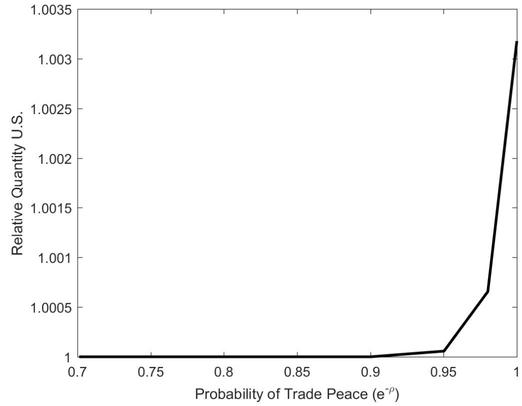
market, causing the wage level there to drop even further to equalize labor supply and demand.

Figure 10 plots aggregate consumption in the U.S. against different probabilities of trade peace, where we normalize the quantity imported when $\Psi = 0.7$ to one. The figure shows that the benefits of a higher likelihood of trade peace accrue in a non-linear fashion. Changing the probability of trade peace from 0.7 to 0.9 has no impact on aggregate consumption, since no imports are made from China for such levels of Ψ. As the probability of trade peace increases further, however, the costs of importing from China fall more and more sharply, leading to higher and higher benefits from switching to China. Aggregate U.S. consumption when the probability of trade peace is 0.95 is only 0.01% higher than when Ψ is 0.7, while at $\Psi = 1$ the benefits exceed 0.3%. This exercise highlights that reducing the likelihood of a trade war only has significant effects if it goes all the way towards ruling out trade wars.

5 Conclusion

This paper analyzes the impact of changes in trade policy on procurement patterns along a supply chain. We develop a theoretical model in which importers' ability to solve a quality control problem depends upon exporters' beliefs about the possibility of

Figure 10: Aggregate Consumption vs. Continuation Probability (Ψ)



a trade war breaking out between the firms' countries. When the probability of trade peace is small, buyers choose American-style procurement, characterized by competitive bidding for large, infrequent orders, and costly inspections to ensure the provision of high-quality goods. When the probability of trade peace is high, buyers can induce sellers to provide high quality without inspections by paying them a premium above their costs over a long-term relationship. We show that changes in trade policy that reduce the likelihood of trade wars increase welfare by lowering procurement costs.

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

Our findings suggest that an important but under-examined aspect of trade agreements in a world with already low tariffs may be their affect 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.[30] The extent to which such

[30]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.

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.

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Appendix

A Analytical Results

A.1 Proof that the Cost Function is Convex

American system

Under the American system, the cost function is given by

$$C_A(x) = \frac{f + \bar{\theta}\frac{\omega}{\Upsilon}x + m}{1 - \delta(x)},$$

The second derivative of the cost function with respect to x is

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

Re-writing this, we obtain

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

Thus, the function is convex if and only if

$$[1 + \delta(x)]\left[x + \frac{f+m}{\bar{\theta}(\omega/\Upsilon)}\right] > 2\frac{1 - \delta(x)}{(r/q)}.$$

Consider the case of $r \to 0$. If x/q is finite, then this condition converges to

$$2\left[x + \frac{f+m}{\bar{\theta}(\omega/\Upsilon)}\right] > 2x,$$

where we have used the expression for $\delta(x) = e^{-(r/q)x}$ and L'Hopital's rule. This expression holds since $(f + m) > 0$.

Japanese system

Define $\tilde{\delta}(x) = e^{-(r+\rho)x/q}$. Under the Japanese system, the cost function is then given by

$$C_J(x) = \frac{f + \underline{\theta}\frac{\omega}{\Upsilon}x + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}x}{1 - \delta(x)}.$$

The second derivative of this cost function with respect to x is

$$
\begin{aligned}
C_J''(x) = & \frac{2\left(\frac{r}{q}\right)^2 \delta(x)^2 \left[f + \underline{\theta}\frac{\omega}{\Upsilon}x + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}x\right]}{\left[1-\delta(x)\right]^3} \\
& + \frac{\left(\frac{r}{q}\right)^2 \delta(x) \left[f + \underline{\theta}\frac{\omega}{\Upsilon}x + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}x\right]}{\left[1-\delta(x)\right]^2} \\
& - \frac{2\left(\frac{r}{q}\right) \delta(x) \left[\underline{\theta}\frac{\omega}{\Upsilon} + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}\left(1 + \left(\frac{r+\rho}{q}\right)x\right)\right]}{\left[1-\delta(x)\right]^2} \\
& + \frac{\left(\frac{r+\rho}{q}\right)^2 \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}x + 2\left(\frac{r+\rho}{q}\right)\frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}}{1-\delta(x)}.
\end{aligned}
$$

Combining terms, we obtain

$$
\begin{aligned}
C_J''(x) = & \frac{\left(\frac{r}{q}\right)^2 \delta(x) \left[f + \left(\underline{\theta} + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\right)\frac{\omega}{\Upsilon}x\right]\left[1+\delta(x)\right]}{\left[1-\delta(x)\right]^3} \\
& - \frac{2\left(\frac{r}{q}\right) \delta(x) \left[\underline{\theta}\frac{\omega}{\Upsilon} + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}\left(1 + \left(\frac{r+\rho}{q}\right)x\right)\right]\left[1-\delta(x)\right]}{\left[1-\delta(x)\right]^3} \\
& + \frac{\left(\frac{r+\rho}{q}\right)\frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}\left[2 + \left(\frac{r+\rho}{q}\right)x\right]\left[1-\delta(x)\right]^2}{\left[1-\delta(x)\right]^3}.
\end{aligned}
$$

Hence, for the cost function to be convex, the numerator of the expression must be greater than zero. This implies the condition

$$
\begin{aligned}
& \delta(x)\left[f + \left(\underline{\theta} + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\right)\frac{\omega}{\Upsilon}x\right]\left[1+\delta(x)\right] \\
& + \left(\frac{r+\rho}{q}\right)\frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}\left[2 + \left(\frac{r+\rho}{q}\right)x\right]\frac{\left[1-\delta(x)\right]^2}{\left(\frac{r}{q}\right)^2} \\
& > 2\delta(x)\left[\underline{\theta}\frac{\omega}{\Upsilon} + \frac{1}{\tilde{\delta}(x)}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}\left(1 + \left(\frac{r+\rho}{q}\right)x\right)\right]\frac{1-\delta(x)}{\left(\frac{r}{q}\right)}.
\end{aligned}
$$

Taking $r \to 0$ and applying L'Hopital's rule, we obtain

$$
\begin{aligned}
& 2\left[f + \left(\underline{\theta} + e^{(\rho/q)x}(\bar{\theta} - \underline{\theta})\right)\frac{\omega}{\Upsilon}x\right] \\
& + \left(\frac{\rho}{q}\right)e^{(\rho/q)x}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}\left[2 + \left(\frac{\rho}{q}\right)x\right]x^2 \\
& > 2\left[\underline{\theta}\frac{\omega}{P}x + e^{(\rho/q)x}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}\left(x + \left(\frac{\rho}{q}\right)x^2\right)\right],
\end{aligned}
$$

which simplifies to

$$2f + \left(\frac{\rho}{q}\right)^2 e^{(\rho/q)x}(\bar{\theta} - \underline{\theta})\frac{\omega}{\Upsilon}x^3 > 0.$$

Since $f > 0$, this condition holds. Therefore, the cost function is convex.

A.2 Proof of Proposition 1

Proof of $x_A^* > x_{FB}^*$

From equation (9) and using the expression for $\delta(x_A^*)$, we have that optimality under the American system requires

$$\frac{\bar{\theta}\frac{\omega}{P}}{1 - \delta(x_A^*)} = \frac{\frac{r}{q}\delta(x_A^*)}{(1 - \delta(x_A^*))^2}\left[f + m + \bar{\theta}\frac{\omega}{P}x_A^*\right]. \tag{A.1}$$

Re-arranging this expression yields

$$\bar{\theta}\frac{\omega}{P} = \frac{\frac{r}{q}\delta(x_A^*)}{1 - \delta(x_A^*)}\left[f + m + \bar{\theta}\frac{\omega}{P}x_A^*\right]. \tag{A.2}$$

Similarly, optimality under the first-best scenario, where $m = 0$, requires

$$\bar{\theta}\frac{\omega}{P} = \frac{\frac{r}{q}\delta(x_{FB}^*)}{1 - \delta(x_{FB}^*)}\left[f + \bar{\theta}\frac{\omega}{P}x_{FB}^*\right]. \tag{A.3}$$

Since the left-hand side of equations (A.2) and (A.3) is the same, we can set them equal and obtain

$$\delta(x_A^*)\frac{f + m + \bar{\theta}\frac{\omega}{P}x_A^*}{1 - \delta(x_A^*)} = \delta(x_{FB}^*)\frac{f + \bar{\theta}\frac{\omega}{P}x_{FB}^*}{1 - \delta(x_{FB}^*)}.$$

From this expression, we obtain the sequence of inequalities:

$$\begin{aligned}
\delta(x_A^*)\frac{f + m + \bar{\theta}\frac{\omega}{P}x_A^*}{1 - \delta(x_A^*)} &= \delta(x_{FB}^*)\frac{f + \bar{\theta}\frac{\omega}{P}x_{FB}^*}{1 - \delta(x_{FB}^*)} \\
&< \delta(x_{FB}^*)\frac{f + \bar{\theta}\frac{\omega}{P}x_A^*}{1 - \delta(x_A^*)} \\
&< \delta(x_{FB}^*)\frac{f + m + \bar{\theta}\frac{\omega}{P}x_A^*}{1 - \delta(x_A^*)},
\end{aligned}$$

where the first inequality follows because the fraction is exactly the cost function $C(q)$ in the first-best case from (7), which is minimized at x_{FB}^*. The second inequality follows since m is a positive constant. Comparing the left-hand side and the right-hand side of the expression yields $\delta(x_A^*) < \delta(x_{FB}^*)$, and therefore $x_A^* > x_{FB}^*$, as claimed.

Proof of $x_J^* < x_{FB}^*$

The proof proceeds along the same lines as in the case of the shipment quantities under the American system. Define $\tilde{\delta}(x_J^*) \equiv e^{-(r+\rho)x_J^*/q}$. From equation (9) and using the expression for $\tilde{\delta}(x_J^*)$ and $\underline{\theta} = 0$, we have that optimality under the Japanese system requires

$$\frac{\frac{\omega}{P}\bar{\theta}\frac{1}{\tilde{\delta}(x_J^*)} + \frac{(r+\rho)/q}{\tilde{\delta}(x_J^*)}\bar{\theta}\frac{\omega}{P}x_J^* - \kappa\Psi'(x_J^*)}{1-\delta(x_J^*)}$$
$$= \frac{\frac{r}{q}\delta(x_J^*)}{(1-\delta(x_J^*))^2}\left[f + \bar{\theta}\frac{\omega}{P}\frac{x_J^*}{\tilde{\delta}(x_J^*)} + \kappa(1-\Psi(x_J^*))\right].$$

This expression can be simplified to

$$\frac{\frac{\frac{\omega}{P}\bar{\theta}}{\tilde{\delta}(x_J^*)}\left[1+\frac{r+\rho}{q}x_J^*\right] - \kappa\Psi'(x_J^*)}{1-\delta(x_J^*)} = \frac{\frac{r}{q}\delta(x_J^*)}{(1-\delta(x_J^*))^2}\left[f + \bar{\theta}\frac{\omega}{P}\frac{x_J^*}{\tilde{\delta}(x_J^*)} + \kappa(1-\Psi(x_J^*))\right]. \tag{A.4}$$

Re-arranging and using the expression for $\Psi'(x)$ yields

$$\bar{\theta}\frac{\omega}{P} = \frac{\frac{r}{q}\delta(x_J^*)}{(1-\delta(x_J^*))}\frac{f\tilde{\delta}(x_J^*) + \frac{\omega}{P}\bar{\theta}x_J^* + \kappa\tilde{\delta}(x_J^*)(1-\Psi(x_J^*))}{1+\frac{r+\rho}{q}x_J^*} - \frac{\kappa\frac{\rho}{q}\delta(x_J^*)}{1+\frac{r+\rho}{q}x_J^*}\Psi(x_J^*). \tag{A.5}$$

As in the case of the American system, we set equation (A.5) equal to the expression for the first-best solution (A.3) and obtain

$$\delta(x_{FB}^*)\frac{f+\frac{\omega}{P}\bar{\theta}x_{FB}^*}{1-\delta(x_{FB}^*)} = \delta(x_J^*)\frac{f\tilde{\delta}(x_J^*) + \frac{\omega}{P}\bar{\theta}x_J^* + \kappa\tilde{\delta}(x_J^*)(1-\Psi(x_J^*))}{(1-\delta(x_J^*))\left(1+\frac{r+\rho}{q}x_J^*\right)}$$
$$-\kappa\frac{\rho}{q}\left[\frac{\tilde{\delta}(x_J^*)}{1+\frac{r+\rho}{q}x_J^*}\Psi(x_J^*)\right]..$$

Using the assumption $\kappa/q \approx 0$, the last term in the previous expression disappears. Dividing both sides by $\tilde{\delta}(x_J^*)$ then yields

$$\begin{aligned}
\frac{\delta(x_{FB}^*)}{\tilde{\delta}(x_J^*)} \frac{f + \frac{\omega}{P}\bar{\theta} x_{FB}^*}{1-\delta(x_{FB}^*)} &= \delta(x_J^*) \frac{f + \frac{\omega}{P}\bar{\theta}\frac{x_J^*}{\tilde{\delta}(x_J^*)} + \kappa(1-\Psi(x_J^*))}{(1-\delta(x_J^*))\left(1 + \frac{r+\rho}{q} x_J^*\right)} \\
&< \delta(x_J^*) \frac{f + \frac{\omega}{P}\bar{\theta}\frac{x_J^*}{\tilde{\delta}(x_J^*)} + \kappa(1-\Psi(x_J^*))}{1-\delta(x_J^*)} \\
&< \delta(x_J^*) \frac{f + \frac{\omega}{P}\bar{\theta}\frac{x_{FB}^*}{\tilde{\delta}(x_{FB}^*)} + \kappa(1-\Psi(x_{FB}^*))}{1-\delta(x_{FB}^*)} \\
&= \frac{\delta(x_J^*)}{\tilde{\delta}(x_{FB}^*)} \frac{f\tilde{\delta}(x_{FB}^*) + \frac{\omega}{P}\bar{\theta} x_{FB}^* + \tilde{\delta}(x_{FB}^*)\kappa(1-\Psi(x_{FB}^*))}{1-\delta(x_{FB}^*)} \\
&< \frac{\delta(x_J^*)}{\tilde{\delta}(x_{FB}^*)} \frac{f + \frac{\omega}{P}\bar{\theta} x_{FB}^* + \kappa(1-\Psi(x_{FB}^*))}{1-\delta(x_{FB}^*)},
\end{aligned}$$

where the first inequality holds because $1 + \frac{r+\rho}{q} x_J^* > 1$, and the second inequality follows from the fact that the fraction is exactly the cost function $C(q)$ in the Japanese system from (7), which is minimized at x_J^*. The final inequality holds because $\tilde{\delta}(x_{FB}^*) < 1$. If κ is sufficiently small, then the term involving κ is negligible and can be disregarded (note that the probability of a trade war multiplying κ is likely also small). In that case, comparing the left-hand side and the right-hand side yields the condition

$$\tilde{\delta}(x_{FB}^*)\delta(x_{FB}^*) < \tilde{\delta}(x_J^*)\delta(x_J^*),$$

from which it follows immediately that $x_J^* < x_{FB}^*$. Thus, if there are no switching costs the quantity ordered under the Japanese system is always smaller than the quantity ordered under the American system. For small switching costs, the relationship still holds, but if κ becomes too large then the desire to save on switching costs outweighs the advantage from ordering more frequently to provide incentives, and firms order less frequently than under first-best. In that case the Japanese system becomes likely the American system.

A.3 Proof of Proposition 2

Consider the case of the American system with $m = 0$. Given per period demand q, the net present value of costs under the optimal order quantity x_A^* is

$$C_A(x_A^*) = \frac{f + \frac{\omega}{P}\bar{\theta} x_A^*}{1-\delta(x_A^*)}.$$

Since this cost function is the same as in the first-best case, the American solution corresponds to the first-best solution and therefore $x_A^* = x_{FB}^*$ and $C_A(x_A^*) = C_{FB}(x_{FB}^*)$.

Consider now the Japanese system with $\rho = 0$, and hence $\Psi = 1$. As demonstrated in the main text, costs are declining in Ψ and thus are lowest under the Japanese system when $\Psi = 1$. Furthermore, assume no switching costs, and hence $\kappa = 0$. The costs under the Japanese system must satisfy

$$C_J(x_J^*) = \frac{f + \frac{\omega}{P}\bar{\theta}\frac{x_J^*}{\delta(x_J^*)}}{1 - \delta(x_J^*)} > \frac{f + \frac{\omega}{P}\bar{\theta}x_J^*}{1 - \delta(x_J^*)} > \frac{f + \frac{\omega}{P}\bar{\theta}x_{FB}^*}{1 - \delta(x_{FB}^*)} = C_{FB}(x_{FB}^*) = C_A(x_A^*), \tag{A.6}$$

where the first inequality holds because $\delta(x_J^*) < 1$ since $r > 0$, and the second inequality holds because x_J^* is not the cost-minimizing batch size in the first-best cost function. Hence, costs under the Japanese system are strictly greater than under the American system. For $\kappa > 0$, costs under the Japanese system are even greater and therefore must also be higher than under the American system.

Since the cost function under the American system is monotonely increasing in m, there must exist $\underline{m}$ such that for $\rho = 0$ the inequalities in (A.6) become an equality:

$$\frac{f + \frac{\omega}{P}\bar{\theta}\frac{x_J^*}{\delta(x_J^*)}}{1 - \delta(x_J^*)} = \frac{f + \frac{\omega}{P}\bar{\theta}x_A^* + \underline{m}}{1 - \delta(x_A^*)}. \tag{A.7}$$

This equation implicitly defines $\underline{m}$. If $\kappa > 0$, the threshold level must increase. Finally, if $m > \underline{m}$, then the left-hand side of equation (A.7) must be strictly smaller than the right-hand side. Since the costs under the Japanese system are increasing in ρ (decreasing in Ψ), and since costs under the Japanese system diverge to infinity as $\rho \to \infty$, for any finite $m > \underline{m}$ there must exist a Ψ^{Switch} such that the costs under both system are equal.

A.4 Proof of Proposition 3

Recall that the probability of trade peace Ψ is inversely related to the arrival rate of trade wars ρ. Define $\tilde{\delta}(x_J^*) \equiv e^{-(r+\rho)x_J^*/q}$. Applying the implicit function theorem to the optimality condition (9) under the Japanese system, we obtain that $dx_J^*/d\rho = A/B$, where the numerator A is given by

$$A = \bar{\theta}\frac{\omega}{P}\frac{x_J^*}{q} + r\frac{x_J^*}{q}\delta(x_J^*)\tilde{\delta}(x_J^*)\frac{\frac{f}{q} + \frac{\kappa}{q} - 2\frac{\kappa}{q}\Psi(x_J^*)}{1 - \delta(x_J^*)} + \left[\frac{q - 2\rho x_J^*}{q}\right]\frac{\kappa}{q}\tilde{\delta}(x_J^*)\Psi(x_J^*), \tag{A.8}$$

and the denominator B equals

$$B = -C + D - E - F + G + H - I, \tag{A.9}$$

where

$$C = \bar{\theta}\frac{\omega}{P}\frac{r+\rho}{q}, \quad D = \frac{\rho(2\rho + r)}{q}\frac{\kappa}{q}\tilde{\delta}(x_J^*)\Psi(x_J^*),$$

$$E = \frac{\left(\frac{r}{q}\right)^2 \delta(x_J^*)\left[f\tilde{\delta}(x_J^*) + \bar{\theta}\frac{\omega}{P}x_J^* + \kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)\right]}{1-\delta(x_J^*)},$$

$$F = \frac{\frac{r}{q}\tilde{\delta}(x_J^*)\delta(x_J^*)(r+\rho)\left(\frac{f+\kappa}{q}\right)}{1-\delta(x_J^*)}, \quad G = \frac{\frac{r}{q}\delta(x_J^*)\bar{\theta}\frac{\omega}{P}}{1-\delta(x_J^*)},$$

$$H = \frac{\frac{r}{q}\delta(x_J^*)\frac{\kappa}{q}(2\rho + r)\tilde{\delta}(x_J^*)\Psi(x_J^*)}{1-\delta(x_J^*)},$$

$$I = \frac{\left(\frac{r}{q}\right)^2 [\delta(x_J^*)]^2\left[f\tilde{\delta}(x_J^*) + \bar{\theta}\frac{\omega}{P}x_J^* + \kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)\right]}{[1-\delta(x_J^*)]^2}.$$

Under the assumption that $\kappa/q \approx 0$, term A simplifies to

$$A = \bar{\theta}\frac{\omega}{P}\frac{x_J^*}{q} + r\frac{x_J^*}{q}\delta(x_J^*)\tilde{\delta}(x_J^*)\frac{\frac{f}{q}}{1-\delta(x_J^*)} > 0.$$

To show that $dx_J^*/d\rho < 0$, it therefore remains to prove that B is negative. With $\kappa/q \approx 0$, terms D and H are approximately zero, and term F becomes

$$F = \frac{\frac{r}{q}\tilde{\delta}(x_J^*)\delta(x_J^*)(r+\rho)\frac{f}{q}}{1-\delta(x_J^*)}.$$

Thus, we need to show that

$$C + E + F + I > G. \tag{A.10}$$

Since r is small and $x < q$, we can use the Taylor approximation $\delta(x_J^*) \approx 1 - \frac{r}{q}x_J^*$. Applying this approximation, we obtain

$$E \approx \frac{\frac{r}{q}\delta(x_J^*)\left[f\tilde{\delta}(x_J^*) + \bar{\theta}\frac{\omega}{P}x_J^*\right]}{x_J^*},$$

where the κ-term disappears because $\kappa/q \approx 0$,

$$F \approx \frac{\tilde{\delta}(x_J^*)\delta(x_J^*)(r+\rho)\frac{f}{q}}{x_J^*},$$

$$G \approx \frac{\bar{\theta}\frac{\omega}{P}}{x_J^*} - \bar{\theta}\frac{\omega}{P}\frac{r}{q},$$

and

$$\begin{aligned}
I \approx & \frac{f\tilde{\delta}(x_J^*) + \bar{\theta}\frac{\omega}{P}x_J^* + \kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)}{(x_J^*)^2} \\
& + \left(\frac{r}{q}\right)^2 \left[f\tilde{\delta}(x_J^*) + \bar{\theta}\frac{\omega}{P}x_J^* + \kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)\right] \\
& - 2\frac{r}{q}\left[\frac{f\tilde{\delta}(x_J^*)}{x_J^*} + \bar{\theta}\frac{\omega}{P}\right].
\end{aligned}$$

Adding together the approximated terms E and F yields

$$E + F = 2\frac{\left(\frac{r}{q}\right)\tilde{\delta}(x_J^*)\delta(x_J^*)f}{x_J^*} + \frac{\left(\frac{\rho}{q}\right)\tilde{\delta}(x_J^*)\delta(x_J^*)f}{x_J^*} + \tfrac{r}{q}\delta(x_J^*)\bar{\theta}\tfrac{\omega}{P}.$$

We can apply another approximation to the first term of this expression to obtain

$$2\frac{\left(\frac{r}{q}\right)\tilde{\delta}(x_J^*)\delta(x_J^*)f}{x_J^*} \approx 2\frac{\left(\frac{r}{q}\right)\tilde{\delta}(x_J^*)f}{x_J^*} - 2\left(\tfrac{r}{q}\right)^2\tilde{\delta}(x_J^*)f.$$

Plugging all these expressions into the condition (A.10) gives

$$\begin{aligned}
\frac{\bar{\theta}\frac{\omega}{P}}{x_J^*} - \bar{\theta}\frac{\omega}{P}\frac{r}{q} < & \bar{\theta}\frac{\omega}{P}\frac{r+\rho}{q} + 2\left(\frac{r}{q}\right)\frac{\tilde{\delta}(x_J^*)f}{x_J^*} - 2\left(\tfrac{r}{q}\right)^2\tilde{\delta}(x_J^*)f \\
& + \frac{\left(\frac{\rho}{q}\right)\tilde{\delta}(x_J^*)\delta(x_J^*)f}{x_J^*} + \tfrac{r}{q}\delta(x_J^*)\bar{\theta}\tfrac{\omega}{P} \\
& + \frac{f\delta(x_J^*) + \kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)}{(x_J^*)^2} + \frac{\bar{\theta}\frac{\omega}{P}}{x_J^*} \\
& + \left(\frac{r}{q}\right)^2\left[\tilde{\delta}(x_J^*)f + \bar{\theta}\tfrac{\omega}{P}x_J^* + \kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)\right] \\
& - 2\left(\frac{r}{q}\right)\left[\frac{\tilde{\delta}(x_J^*)f}{x_J^*} + \bar{\theta}\tfrac{\omega}{P}\right].
\end{aligned}$$

Cancelling terms, the expression simplifies to

$$0 < \bar{\theta}\frac{\omega}{P}\frac{\rho}{q} - \left(\frac{r}{q}\right)^2 \tilde{\delta}(x_J^*)f + \frac{\left(\frac{\rho}{q}\right)\tilde{\delta}(x_J^*)\delta(x_J^*)f}{x_J^*} + \tfrac{r}{q}\delta(x_J^*)\bar{\theta}\tfrac{\omega}{P}$$
$$+ \frac{1}{(x_J^*)^2}\tilde{\delta}(x_J^*)f + \frac{\kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)}{(x_J^*)^2} + \left(\frac{r}{q}\right)^2 \left[\bar{\theta}\tfrac{\omega}{P}x_J^* + \kappa(1-\tilde{\delta}(x_J^*))\Psi(x_J^*)\right].$$

Note that only the second term in this expression is negative. A sufficient condition for it to hold is

$$\frac{1}{(x_J^*)^2} > \frac{r^2}{q^2}.$$

Since $r < 1$and $q > x_J^*$, this condition is satisfied and hence $dx_J^*/d\rho < 0$, as claimed.

A.5 Proof of Proposition 4

1a. Proof of $dx_A^*/df > 0$

From the implicit function theorem, we have that

$$\frac{dx_A^*}{df} = \frac{\frac{r}{q}\delta(x_A^*)}{B},$$

where

$$\begin{aligned} B = & \left(\frac{r}{q}\right)^2 \delta(x_A^*)\left[f + m + \tfrac{\omega}{P}\bar{\theta}x_A^*\right] \\ & - \left(\frac{r}{q}\right)\delta(x_A^*)\tfrac{\omega}{P}\bar{\theta} \\ & + \frac{1}{1-\delta(x_A^*)}\left(\frac{r}{q}\right)^2 [\delta(x_A^*)]^2 \left[f + m + \tfrac{\omega}{P}\bar{\theta}x_A^*\right]. \end{aligned}$$

Dividing B by the numerator $\frac{r}{q}\delta(x_A^*)$ and combining the first and the last term in B yields $dx_A^*/df = 1/C$, where

$$C = \frac{1}{1-\delta(x_A^*)}\left(\frac{r}{q}\right)\left[f + m + \tfrac{\omega}{P}\bar{\theta}x_A^* - (1-\delta(x_A^*))\tfrac{q}{r}\tfrac{\omega}{P}\bar{\theta}\right].$$

Using the approximation $\delta(x_A^*) \approx 1 - \frac{r}{q}x_A^*$, the term in parentheses can be simplified to yield

$$C = \frac{1}{1-\delta(x_A^*)}\left(\frac{r}{q}\right)[f + m].$$

Therefore, we obtain

$$\frac{dx_A^*}{df} = \frac{1}{\frac{1}{1-\delta(x_A^*)}\left(\frac{r}{q}\right)[f+m]} > 0. \tag{A.11}$$

1b. Proof of $dx_J^*/df > 0$

Using the implicit function theorem and equation (9), we obtain that

$$\frac{dx_J^*}{df} = \frac{\tilde{\delta}(x_J^*)}{B},$$

where

$$\begin{aligned} B = & \left(\frac{r+\rho}{q}\right)\tfrac{\omega}{P}\bar{\theta}\left(\frac{q}{r}\right)\frac{1-\tilde{\delta}(x_J^*)}{\tilde{\delta}(x_J^*)} - \rho\frac{\kappa}{q}\Psi(x)\delta(x)\left(\frac{2\rho+r}{r}\right)\frac{1-\tilde{\delta}(x_J^*)}{\tilde{\delta}(x_J^*)} \\ & + \frac{r}{q}\frac{1}{1-\tilde{\delta}(x_J^*)}\left[f\delta(x_J^*) + \tfrac{\omega}{P}\bar{\theta}x_J^* + \kappa\left(1-\Psi(x_J^*)\right)\delta(x_J^*)\right] \\ & + f\left(\frac{r+\rho}{q}\right)\delta(x_J^*) - \tfrac{\omega}{P}\bar{\theta} + \frac{\kappa}{q}\delta(x_J^*)\left[\rho+r-(2\rho+r)\Psi(x_J^*)\right]. \end{aligned}$$

Using the assumption that $\kappa/q \approx 0$, the expression simplifies to

$$\begin{aligned} B = & \left(\frac{r+\rho}{q}\right)\tfrac{\omega}{P}\bar{\theta}\left(\frac{q}{r}\right)\frac{1-\tilde{\delta}(x_J^*)}{\tilde{\delta}(x_J^*)} + f\left(\frac{r+\rho}{q}\right)\delta(x_J^*) - \tfrac{\omega}{P}\bar{\theta} \\ & + \frac{r}{q}\frac{1}{1-\tilde{\delta}(x_J^*)}\left[f\delta(x_J^*) + \tfrac{\omega}{P}\bar{\theta}x_J^* + \kappa\left(1-\Psi(x_J^*)\right)\delta(x_J^*)\right]. \end{aligned}$$

Using the approximation $\tilde{\delta}(x_J^*) \approx 1 - \frac{r}{q}x_J^*$, the expression becomes

$$\begin{aligned} B = & \left(\frac{r+\rho}{q}\right)\tfrac{\omega}{P}\bar{\theta}\left(\frac{q}{r}\right)\frac{1-\tilde{\delta}(x_J^*)}{\tilde{\delta}(x_J^*)} + f\left(\frac{r+\rho}{q}\right)\delta(x_J^*) + f\frac{\delta(x_J^*)}{x_J^*} \\ & + \kappa\left(1-\Psi(x_J^*)\right)\frac{\delta(x_J^*)}{x_J^*} > 0. \end{aligned}$$

Thus, both the numerator and the denominator are positive, and therefore $dx_J^*/df > 0$, as claimed.

2. Proof of $dx_A^*/dm > 0$

Since f and m appear in equation (9) in the same way, the proof follows exactly the same steps as the proof that $dx_A^*/df > 0$. Thus, $dx_A^*/dm > 0$.

3a. Proof of $dx_A^*/d\bar{\theta} < 0$

Applying the implicit function theorem to equation (9), we obtain

$$\frac{dx_A^*}{d\bar{\theta}} = \frac{1 - \frac{\frac{r}{q}\delta(x_A^*)}{1-\delta(x_A^*)}x_A^*}{C},$$

where

$$C = -\frac{\left(\frac{r}{q}\right)^2 \delta(x_A^*)}{[1-\delta(x_A^*)]^2}\left[f + \frac{\omega}{P}\bar{\theta}x_A^* + m\right] + \frac{\left(\frac{r}{q}\right)\delta(x_A^*)}{[1-\delta(x_A^*)]}\frac{\omega}{P}\bar{\theta}.$$

Dividing both the numerator and the denominator by $\left(\frac{r}{q}\right)\delta(x_A^*)/\left[1-\delta(x_A^*)\right]$ yields

$$\frac{dx_A^*}{d\bar{\theta}} = \frac{\frac{1-\delta(x_A^*)}{\frac{r}{q}\delta(x_A^*)} - x_A^*}{-\frac{(r/q)}{1-\delta(x_A^*)}\left[f + \frac{\omega}{P}\bar{\theta}x_A^* + m\right] + \frac{\omega}{P}\bar{\theta}}.$$

Using the approximation $\delta(x_A^*) \approx 1 - \frac{r}{q}x_A^*$ to the $1-\delta(x_A^*)$ terms, we obtain

$$\frac{dx_A^*}{d\bar{\theta}} \approx \frac{\frac{x_A^*}{\delta(x_A^*)} - x_A^*}{-\frac{f+m}{x_A^*}} = -\frac{\left(1-\delta(x_A^*)\right)\left(x_A^*\right)^2}{\left(f+m\right)\delta(x_A^*)} < 0,$$

as required.

3b. Proof of $dx_J^*/d\bar{\theta} < 0$

Applying the implicit function theorem to equation (9) yields

$$\frac{dx_J^*}{d\bar{\theta}} = \frac{1 + \left(\frac{r+\rho}{q}\right)x_J^* - \frac{\frac{r}{q}\delta(x_J^*)}{1-\delta(x_J^*)}x_J^*}{D},$$

where

$$\begin{aligned}D = &-\left(\frac{r+\rho}{q}\right)\frac{\omega}{P}\bar{\theta} + \rho\frac{\kappa}{q}\Psi(x)\tilde{\delta}(x)\left(\frac{2\rho+r}{q}\right) \\ &- \frac{\left(\frac{r}{q}\right)^2\delta(x_J^*)}{[1-\delta(x_J^*)]^2}\left[f\tilde{\delta}(x_J^*) + \tfrac{\omega}{P}\bar{\theta}x_J^* + \kappa(1-\Psi(x_J^*))\tilde{\delta}(x_J^*)\right] \\ &- \frac{\left(\frac{r}{q}\right)\delta(x_J^*)}{[1-\delta(x_J^*)]}\left[f\left(\frac{r+\rho}{q}\right)\tilde{\delta}(x_J^*) - \tfrac{\omega}{P}\bar{\theta}x_J^*\right] \\ &- \frac{\left(\frac{r}{q}\right)\delta(x_J^*)}{[1-\delta(x_J^*)]}\frac{\kappa}{q}\tilde{\delta}(x_J^*)\left[\rho + r - (2\rho+r)\,\Psi(x_J^*)\right].\end{aligned}$$

Using the fact that $\kappa/q \approx 0$ and multiplying both the numerator and the denominator by $(r/q)\delta(x_J^*)/1-\delta(x_J^*)$, we obtain

$$\frac{dx_J^*}{d\bar{\theta}} = \frac{\left[1+\left(\frac{r+\rho}{q}\right)x_J^*\right]\frac{1-\delta(x_J^*)}{\frac{r}{q}\delta(x_J^*)} - x_J^*}{E},$$

where

$$\begin{aligned}E = &-\left(\frac{r+\rho}{q}\right)\frac{\omega}{P}\bar{\theta}\frac{1-\delta(x_J^*)}{\frac{r}{q}\delta(x_J^*)} - f\left(\frac{r+\rho}{q}\right)\tilde{\delta}(x_J^*) + \tfrac{\omega}{P}\bar{\theta}x_J^* \\ &- \frac{\left(\frac{r}{q}\right)}{[1-\delta(x_J^*)]}\left[f\tilde{\delta}(x_J^*) + \tfrac{\omega}{P}\bar{\theta}x_J^* + \kappa(1-\Psi(x_J^*))\tilde{\delta}(x_J^*)\right].\end{aligned}$$

We apply the approximation $\delta(x_A^*) \approx 1 - \frac{r}{q}x_A^*$ to the $1-\delta(x_A^*)$ terms to obtain, for the numerator,

$$\left[1+\left(\tfrac{r+\rho}{q}\right)x_J^*\right]\frac{1-\delta(x_J^*)}{\frac{r}{q}\delta(x_J^*)} - x_J^* \approx \left[1+\left(\tfrac{r+\rho}{q}\right)x_J^*\right]\frac{x_J^*}{\delta(x_J^*)} - x_J^* > 0,$$

where the inequality holds because $\tilde{\delta}(x_J^*) < 1$. For the denominator we obtain

$$\begin{aligned}E \approx &-\left(\frac{r+\rho}{q}\right)\frac{\omega}{P}\bar{\theta}\frac{x_J^*}{\delta(x_J^*)} - f\left(\frac{r+\rho}{q}\right)\tilde{\delta}(x_J^*) \\ &- f\frac{\tilde{\delta}(x_J^*)}{x_J^*} - \kappa(1-\Psi(x_J^*))\frac{\tilde{\delta}(x_J^*)}{x_J^*} < 0.\end{aligned}$$

Therefore, $dx_J^*/d\bar{\theta} < 0$, as claimed.

4a. Proof of $dx_A^*/dq > 0$ and $d(x_A^*/q)/dq < 0$

From (9), the first derivative of the cost function under the American system is

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

Consider two quantities, $q_1 < q_2$. Suppose we are in the equilibrium of q_1, with $x_A^*(q_1)$ setting the first derivative equal to zero under q_1. If we can show that under q_2 the first derivative of the cost function $C_A'(x)$ becomes negative at $x_A^*(q_1)$, then by the fact that the cost function is strictly concave (established in Proposition A.1) it must be the case that we are to the left of the new cost-minimizing order size, and hence x has to be increased to $x_A^*(q_2) > x_A^*(q_1)$ to attain the new optimum. Thus, we want to show that

$$\frac{\bar{\theta}\frac{w}{\Upsilon}}{1 - e^{-rx_A^*(q_1)/q_2}} - \frac{\left(\frac{r}{q_2}\right) e^{-rx_A^*(q_1)/q_2}\left[f + \bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) + m\right]}{\left[1 - e^{-rx_A^*(q_1)/q_2}\right]^2} < 0.$$

To show this fact, we take the derivative of $C_A'(x)$ with respect to q at the point $x_A^*(q_1)$. This derivative is equal to

$$\begin{aligned}\frac{dC_A'(x)}{dq}\Big|_{x=x_A^*(q_1)} = {} & \frac{\bar{\theta}\frac{w}{\Upsilon}\left(\frac{r}{q^2}\right)x_A^*(q_1)e^{-rx_A^*(q_1)/q}}{\left[1 - e^{-rx_A^*(q_1)/q}\right]^2} \\ & - \frac{2\left(\frac{r^2}{q^3}x_A^*(q_1)\right)e^{-rx_A^*(q_1)/q}\left[f + \bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) + m\right]}{\left[1 - e^{-rx_A^*(q_1)/q}\right]^3} \\ & + \frac{\left[\left(\frac{r}{q^2}\right)e^{-rx_A^*(q_1)/q_2} - \left(\frac{r^2}{q^3}x_A^*(q_1)\right)e^{-rx_A^*(q_1)/q}\right]\left[f + \bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) + m\right]}{\left[1 - e^{-rx_A^*(q_1)/q}\right]^2}.\end{aligned}$$

This derivative is negative if and only if

$$\begin{aligned}&\bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1)\left[1 - e^{-rx_A^*(q_1)/q}\right] - 2\left(\frac{rx_A^*(q_1)}{q}\right)\left[f + \bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) + m\right] \\ &+ \left[1 - \left(\frac{rx_A^*(q_1)}{q}\right)\right]\left[f + \bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) + m\right]\left[1 - e^{-rx_A^*(q_1)/q}\right] < 0.\end{aligned}$$

Since r is close to zero, we can approximate $e^{-rx_A^*(q_1)/q} \approx 1 - rx_A^*(q_1)/q$, which yields

$$\bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) - 2\left[f + \bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) + m\right] + \left[1 - \left(\frac{rx_A^*(q_1)}{q}\right)\right]\left[f + \bar{\theta}\frac{w}{\Upsilon}x_A^*(q_1) + m\right] < 0.$$

This expression holds if and only if

$$-(f+m)-\left(\frac{rx_A^*(q_1)}{q}\right)(f+\bar{\theta}\tfrac{w}{\Upsilon}x_A^*(q_1)+m)<0,$$

which is satisfied. Hence, the derivative of $C_A'(x)$ with respect to q is negative at the initial optimization point, and thus x has to increase to the new equilibrium. Thus, $dx_A^*/dq>0$.

To see that $d(x_A^*/q)/dq<0$, define $\psi_A=x_A/q$. Then, from the first order condition (9), we have

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

From the implicit function theorem, we find that

$$\frac{d\psi_A}{dq}=-\frac{\left(\frac{q}{r}\right)[f+m]}{f+m+\bar{\theta}\frac{w}{\Upsilon}q\psi_A}<0,$$

and hence the time between shipments decreases.

4b. Proof of $dx_J^*/dq>0$ and $d(x_J^*/q)/dq<0$

From (9), the first derivative of the cost function under the Japanese system is

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

We proceed as in the case of the American system. Assume that we are at the optimum for some quantity q_1. We want to show that the derivative of the cost function for $q_2>q_1$ at $x^*(q_1)$ is negative, and hence the order size has to be increased to the new optimum. We take the derivative of $C_J'(x)$ with respect to q at the point $x_J^*(q_1)$. This derivative is equal to

$$\frac{dC'_J(x)}{dq}|_{x=x^*_J(q_1)} = \frac{-\frac{(r+\rho)}{q^2}x^*_J(q_1)\bar{\theta}\frac{w}{\Upsilon}e^{(r+\rho)x^*_J(q_1)/q}\left[2+(r+\rho)\frac{x^*_J(q_1)}{q_1}\right]}{1-e^{-rx^*_J(q_1)/q}}$$
$$+\frac{\left(\frac{r}{q^2}\right)e^{-rx^*_J(q_1)/q}f\left[1-\frac{rx^*_J(q_1)}{q}\right]}{\left[1-e^{-rx^*_J(q_1)/q}\right]^2}-\frac{2\left(\frac{r^2}{q^3}\right)x^*_J(q_1)\left(e^{-rx^*_J(q_1)/q}\right)^2 f}{\left[1-e^{-rx^*_J(q_1)/q}\right]^3}$$
$$+\frac{\left(\frac{r}{q^2}\right)e^{-rx^*_J(q_1)/q}\bar{\theta}\frac{w}{\Upsilon}e^{(r+\rho)x^*_J(q_1)/q}x^*_J(q_1)\left[2+2(r+\rho)\frac{x^*_J(q_1)}{q}-r\frac{x^*_J(q_1)}{q}\right]}{\left[1-e^{-rx^*_J(q_1)/q}\right]^2}$$
$$-2\frac{\left(\frac{r^2}{q^3}\right)\left(e^{-rx^*_J(q_1)/q}\right)^2\bar{\theta}\frac{w}{\Upsilon}e^{(r+\rho)x^*_J(q_1)/q}\left(x^*_J(q_1)\right)^2}{\left[1-e^{-rx^*_J(q_1)/q}\right]^3}.$$

This derivative is less than zero if and only if

$$\begin{aligned}
&-\tfrac{(r+\rho)}{q^2}x^*_J(q_1)\bar{\theta}\tfrac{w}{\Upsilon}e^{(r+\rho)x^*_J(q_1)/q}\left[2+(r+\rho)\tfrac{x^*_J(q_1)}{q_1}\right]\left[1-e^{-rx^*_J(q_1)/q}\right]^2+\\
&+\left(\tfrac{r}{q^2}\right)e^{-rx^*_J(q_1)/q}f\left\{\left[1-\tfrac{rx^*_J(q_1)}{q}\right]\left[1-e^{-rx^*_J(q_1)/q}\right]-2\left(\frac{r}{q}\right)x^*_J(q_1)e^{-rx^*_J(q_1)/q}\right\}\\
&+\left(\tfrac{r}{q^2}\right)e^{-rx^*_J(q_1)/q}\bar{\theta}\tfrac{w}{\Upsilon}e^{(r+\rho)x^*_J(q_1)/q}x^*_J(q_1)\Big[\left[2\left(1+(r+\rho)\tfrac{x^*_J(q_1)}{q}\right)-r\tfrac{x^*_J(q_1)}{q}\right]\left[1-e^{-rx^*_J(q_1)/q}\right]\\
&-2r\tfrac{x^*_J(q_1)}{q}e^{-rx^*_J(q_1)/q}\Big]<0.
\end{aligned} \tag{A.12}$$

Using the fact that r is small and the approximation $e^{-rx/q}\approx 1-rx/q$, the second term in this expression can be simplified to

$$\begin{aligned}
&\left(\tfrac{r}{q^2}\right)e^{-rx^*_J(q_1)/q}f\left\{\left[1-\tfrac{rx^*_J(q_1)}{q}\right]\left[1-e^{-rx^*_J(q_1)/q}\right]-2\left(\frac{r}{q}\right)x^*_J(q_1)e^{-rx^*_J(q_1)/q}\right\}\\
\approx&-\left(\tfrac{r}{q^2}\right)e^{-rx^*_J(q_1)/q}f\left(\frac{r}{q}\right)x^*_J(q_1)\left\{\left[1-\tfrac{rx^*_J(q_1)}{q}\right]\right\}<0.
\end{aligned}$$

Therefore, if we can show that the remaining terms in expression (A.12) are negative, the overall derivative will be negative as well. Using the approximation for r small, the sum of these two terms is negative if and only if

$$\begin{aligned}
&=-\tfrac{(r+\rho)}{q^2}x^*_J(q_1)\left[2+(r+\rho)\tfrac{x^*_J(q_1)}{q_1}\right]\left(r^2\tfrac{(x^*_J(q_1))^2}{q^2}\right)\\
&+\left(\tfrac{r}{q^2}\right)e^{-rx^*_J(q_1)/q}\left\{2\left(1+(r+\rho)\tfrac{x^*_J(q_1)}{q}\right)-1-e^{-rx^*_J(q_1)/q}\right\}\left(r\tfrac{(x^*_J(q_1))}{q}\right)<0.
\end{aligned}$$

This can be simplified further to read

$$
\begin{aligned}
& -(r+\rho)\left[2+(r+\rho)\tfrac{x_J^*(q_1)}{q_1}\right]+e^{-rx_J^*(q_1)/q}\left[2r+\rho\right] \\
\approx & -\rho-(r+\rho)^2\tfrac{x_J^*(q_1)}{q_1}-\tfrac{rx_J^*(q_1)}{q}\left[2r+\rho\right]<0.
\end{aligned}
$$

This condition is satisfied, and hence $x^*(q_2) > x^*(q_1)$.

To see that $d(x_J^*/q)/dq < 0$, define $\psi_J = x_J/q$. From (9), the first-order condition under the Japanese system can then be written as

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

From the Implicit Function Theorem, we find that

$$
\frac{d\psi_J}{dq}=-\frac{f\left(\frac{r}{q^2}\right)e^{-r\psi_J}}{A},
$$

where

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

This term can be re-arranged to yield

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

Thus, $d(x_J^*/q)/dq < 0$ as claimed.

B Additional Tables and Figures

B.1 Tables

Table A.1: Simulation parameters

	Baseline
Probability of trade peace ($p = e^{-\rho}$)	0.95
Order quantity (q)	10
Interest rate (r)	0.05
Low, high quality ($\underline{\theta}, \bar{\theta}$)	(0, 10)
Seller fixed cost (f)	0.2
Buyer inspection cost (m)	1
Switching cost (κ)	1

Table A.2: Rauch Regressions for $t = 5$ (Conservative Classification)

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	ln(VPS)	ln(WBS)	ln(Length)	ln(VPS)	ln(WBS)	ln(Length)
d_h^D	-0.192^{***}	-0.182^{***}	0.044^{***}	-0.129^{***}	-0.130^{***}	0.128^{***}
	(0.003)	(0.004)	(0.006)	(0.003)	(0.004)	(0.006)
d_h^{Ref}	-0.119^{***}	-0.109^{***}	0.045^{***}	-0.090^{***}	-0.084^{***}	0.085^{***}
	(0.004)	(0.005)	(0.007)	(0.003)	(0.005)	(0.007)
$\ln(NSupp_{mhcz})$				-0.455^{***}	-0.373^{***}	-0.608^{***}
				(0.001)	(0.001)	(0.002)
Observations	1,556,000	1,556,000	1,556,000	1,556,000	1,556,000	1,556,000
R-Squared	0.874	0.700	0.532	0.904	0.727	0.517
Fixed Effects	mc, z	mc, z	mc, z	mc, z	mc, z	mc, z

Notes: Superscripts *, **, and *** indicate statistical significance at the 10, 5, and 1 percent levels, respectively. Number of observations has been rounded to the nearest 1000 as per U.S. Census Bureau Disclosure Guidelines.

Table A.3: Classification regressions at the importer level, for $t = 15$

	(1)	(2)	(3)	(4)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
d^A_{mhcz}	0.761***	0.847***	−0.114***	−2.261***
	(0.004)	(0.004)	(0.007)	(0.005)
$\ln(Qty_{mhcz})$	0.732***	−0.254***	−0.314***	−0.047***
	(0.001)	(0.001)	(0.002)	(0.001)
Observations	158,000	158,000	158,000	158,000
R-Squared	0.960	0.728	0.869	0.781
Fixed Effects	hc, z	hc, z	hc, z	hc, z

	(5)	(6)	(7)	(8)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
$\ln(SPS_{mhcz})$	0.293***	0.315***	−0.031***	−0.806***
	(0.001)	(0.001)	(0.001)	(0.001)
$\ln(Qty_{mhcz})$	0.744***	−0.253***	−0.334***	−0.067***
	(0.000)	(0.001)	(0.000)	(0.000)
Observations	927,000	927,000	927,000	927,000
R-Squared	0.952	0.593	0.844	0.637
Fixed Effects	hc, z	hc, z	hc, z	hc, z

Notes: Superscripts *, **, and *** indicate statistical significance at the 10, 5, and 1 percent levels, respectively. Number of observations has been rounded to the nearest 1000 as per U.S. Census Bureau Disclosure Guidelines.

Table A.4: Classification regressions at the importer-exporter level, for $t = 15$

	(1)	(2)	(3)	(4)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
d^A_{mhcz}	0.427***	1.307***	−0.102***	−0.195***
	(0.010)	(0.018)	(0.010)	(0.014)
$\ln(Qty_{mxhcz})$	0.555***	−0.208***	−0.108***	0.208***
	(0.002)	(0.002)	(0.002)	(0.003)
Observations	72,000	72, 000	72, 000	72, 000
R-Squared	0.982	0.789	0.970	0.731
Fixed Effects	xhc, z	xhc, z	xhc, z	xhc, z

	(5)	(6)	(7)	(8)
Dependent Variable	ln(QPS)	ln(WBS)	ln(Price)	ln(Length)
SPS_{mhcz}	0.246***	0.531***	−0.056***	−0.121***
	(0.001)	(0.002)	(0.001)	(0.001)
$\ln(Qty_{mxhcz})$	0.637***	−0.142***	−0.140***	0.163***
	(0.000)	(0.001)	(0.001)	(0.001)
Observations	1, 325, 000	1, 325, 000	1, 325, 000	1, 325, 000
R-Squared	0.979	0.652	0.959	0.623
Fixed Effects	xhc, z	xhc, z	xhc, z	xhc, z

Notes: Superscripts *, **, and *** indicate statistical significance at the 10, 5, and 1 percent levels, respectively. Number of observations has been rounded to the nearest 1000 as per U.S. Census Bureau Disclosure Guidelines.