Supplemental Appendix

J Additional Proofs

J.1 Second Order Conditions Hold

American System The second derivative of the average cost yields

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

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

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

Consider the case when $f+m=0$. If the condition holds for this case, it must also hold for $f+m>0$,because (S.1) is increasing in $f+m$. Define $y \equiv rx/q$. Note that for $y=0$ and $f+m=0$ the left-hand side of equation (S.1) is equal to zero. Taking the derivative of the left-hand side of equation (S.1) with respect to y we obtain $1-e^{-y}(1-y)>0$. Thus, the left-hand side of (S.1) is strictly increasing in y for $0<y<1$. Therefore, if $0<y<1$, then $AC_A''(x,q)>0$.

Japanese System

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

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

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

Define $y \equiv rx/q$. For the first term note that $(1 + e^{-y})\, y - 2\,(1 - e^{-y}) > 0$ for $0 < y < 1$. For the second term to be positive, we require that $\left([1 - e^{-y}]\left[2 + y + \left(\frac{\rho}{q}\right) x\right] - 2ye^{-y}\right) > 0$. If $\rho = 0$, then $(\cdot) > 0$ for $0 < y < 1$. Because $(\cdot)$increases in ρ, it must be true that $(\cdot) > 0$ for $\rho > 0$ and $0 < y < 1$. Therefore, if $\rho > 0$ and $0 < y < 1$, then $AC''_J(x) > 0$.

J.2 Continued Proof of Lemma 5.2: Average cost curves are convex and reach a limit

Part 1: Average cost curves are convex

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

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

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

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

which is positive if

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

This expression holds if

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

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

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

Part 2: Average cost curves reach a limit

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

For the first term we have that

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

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

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

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

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

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

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

K Additional Summary Statistics

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

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

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

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

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

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

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

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

Table S.2: Attributes of $mhcz$ Quadruples

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

Source: LFTTD and authors' calculations. Table reports the mean and standard deviation across importer (m) by country (c) by ten-digit Harmonized System category (h) by mode of transport (z) quadruples during our 1992 to 2016 sample period. Observations are based on the cleaned data described in Appendix B. Import values are in real 2009 dollars. The first two columns restrict to our baseline sample of quadruples with at least five transactions, analogous to Table 1. The final two columns restrict to the broader sample of quadruples with two or more transactions. Observation counts are rounded to the nearest thousand per U.S. Census Bureau disclosure guidelines.

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

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

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

L Supplemental A vs J Classification Regressions

Differentiated Products Versus Commodities: We examine whether buyers are more likely to use J procurement for differentiated goods. If differentiated products have higher inspection costs, then by Proposition 2.1 buyers are more likely to use J procurement for them, which implies smaller shipment size, greater frequency, and higher unit import values than products sourced under the A system (Proposition 2.3). Moreover, as discussed in Section 3.3, this J sourcing of differentiated products should be associated with fewer suppliers and longer relationships. We examine these features of the model using the commonly cited measure of product-differentiation from Rauch (1999) in the following $mhcz$-level OLS specification,

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

We consider four dependent variables. The first is the average number of weeks between shipments WBS_{mhcz} as in the main text. We do not consider quantity per shipment or unit value here since the regression compares shipping systems across products, which are recorded in different units.[58] Instead, we use as our second dependent variable the average transaction value per shipment, VPS_{mhcz}, as a measure of average transaction size. Third, we consider the average relationship length ($length_{mhcz}$) as in Section 3.3. Finally, the fourth variable is a measure of the buyer's procurement type, sellers per shipment (SPS_{mhcz}) introduced in the main text. On the right-hand side, $Diff_h$ is a dummy variable indicating that product h is either differentiated or has a reference price, as opposed to being a commodity, according to the product categorization scheme proposed by Rauch (1999).[59] Because the right-hand-side variable of interest varies only at the product level, we are unable to include product fixed effects, so comparisons are made within country-mode bins by including fixed effects at that level (λ_{cz}). Since we cannot standardize quantities to be consistent across products, we control for potential scale effects using value per week (VPW_{mhcz}), rather than quantity per week, which was used in the main text.

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

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

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

Results, reported in Table S.4, are consistent with the model's predictions regarding inspection costs, while providing further support for the use of sellers per shipment to identify buyer types. As indicated in the first three columns of the table, we find that differentiated products are more J: they are shipped with fewer weeks between shipments, the average transaction size is smaller, and the average relationship length is longer. Results in the final column provide further support for this view, as buyer quadruples encompassing differentiated goods tend to have lower sellers per shipment.

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

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

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

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

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

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

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

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

	(1)	(2)	(3)	(4)
Dep. var.	log(WBS_{mhcz})	log(VPS_{mhcz})	log($length_{mhcz}$)	log(SPS_{mhcz})
$Diff_h$	-0.234***	-0.225***	0.073**	-0.082***
	0.026	0.025	0.028	0.025
log(VPW_{mhcz})	-0.464***	0.557***	-0.045***	-0.203***
	0.002	0.002	0.001	0.001
Observations	2,589,000	2,589,000	2,589,000	2,589,000
R-squared	0.611	0.730	0.193	0.278
Fixed effects	cz	cz	cz	cz
Controls	beg, end	beg, end	beg, end	beg, end

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

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

	(1)	(2)	(3)	(4)
Dep. var.	ln(QPS_{mhcz})	ln(WBS_{mhcz})	ln(UV_{mhcz})	ln($length_{mhcz}$)
ln(SPS_{mhcz})	0.413***	0.455***	−0.106**	−0.692***
	0.021	0.022	0.041	0.017
log(QPW_{mhcz})	0.704***	−0.305***	−0.283***	−0.190***
	0.031	0.032	0.019	0.014
Observations	25,500	25,500	25,500	25,500
Fixed effects	hcz	hcz	hcz	hcz
R-squared	0.972	0.756	0.925	0.562
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, and average relationship length. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

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

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	0.500^{***}	0.538^{***}	-0.181^{***}	-0.540^{***}
	0.014	0.014	0.022	0.012
$\log(QPW_{mhcz})$	0.769^{***}	-0.238^{***}	-0.367^{***}	-0.131^{***}
	0.018	0.018	0.022	0.008
Observations	560,000	560,000	560,000	560,000
Fixed effects	hcz	hcz	hcz	hcz
R-squared	0.950	0.712	0.816	0.434
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, and average relationship length. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

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

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	0.443^{***}	0.475^{***}	-0.181^{***}	-0.571^{***}
	0.015	0.015	0.013	0.020
$\log(QPW_{mhcz})$	0.682^{***}	-0.328^{***}	-0.281^{***}	-0.167^{***}
	0.012	0.012	0.017	0.007
Observations	1,215,000	1,215,000	1,215,000	1,215,000
Fixed effects	hcz	hcz	hcz	hcz
R-squared	0.945	0.708	0.856	0.469
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, and average relationship length. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

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

	(1)	(2)	(3)	(4)
Dep. var.	$\ln(QPS_{mhcz})$	$\ln(WBS_{mhcz})$	$\ln(UV_{mhcz})$	$\ln(length_{mhcz})$
$\ln(SPS_{mhcz})$	0.424***	0.458***	−0.120***	−0.556***
	0.030	0.031	0.023	0.022
$\log(QPW_{mhcz})$	0.643***	−0.366***	−0.195***	−0.115***
	0.007	0.007	0.012	0.008
Observations	525,000	525,000	525,000	525,000
Fixed effects	*hcz*	*hcz*	*hcz*	*hcz*
R-squared	0.945	0.708	0.856	0.955
Controls	beg, end	beg, end	beg, end	beg, end

Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of importer by product by country by mode of transport ($mhcz$) bins on bins' sellers per shipment (SPS_{mhcz}) and total quantity shipped per week (QPW_{mhcz}). Industries are assigned using the main 6-digit NAICS industry of the importer based on total employment. QPS_{mhcz}, WBS_{mhcz}, UV_{mhcz}, and $length_{mhcz}$ are average quantity per shipment, average weeks between shipment, average unit value, and average relationship length. All regressions include product by country by mode of transport (hcz) fixed effects, control for the beginning and end week of the quadruple, and exclude quadruples with less than five shipments. Standard errors, adjusted for clustering by country (c) and product (h) are reported below coefficient estimates. ***, **, and * represent statistical significance at the 1, 5 and 10 percent levels.

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

	(1)	(2)	(3)	(4)
Dep. var.	$\log(QPS_{mxhcz})$	$\ln(WBS_{mxhcz})$	$\ln(UV_{mxhcz})$	$\ln(length_{mxhcz})$
$\ln(SPS_{mhcz})$	0.100***	0.696***	−0.062***	−0.302***
	0.015	0.041	0.006	0.011
$\ln(QPW_{mxhcz})$	0.511***	−0.171***	−0.130***	−0.241***
	0.010	0.009	0.011	0.008
Observations	4,783,000	4,783,000	4,783,000	4,783,000
R-squared	0.966	0.621	0.953	0.786
Fixed effects	*xhcz*	*xhcz*	*xhcz*	*xhcz*
Controls	beg, end	beg, end	beg, end	beg, end

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

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

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

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

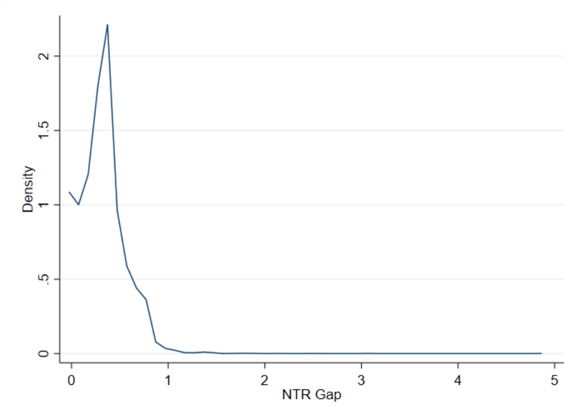
M Description of PNTR

This section provides more detail on the US granting permanent normal trade relations (PNTR) to China. US imports from non-market economies such as China are generally subject to relatively high "column two" tariff rates originally set under the Smoot-Hawley Tariff Act of 1930, as opposed to the generally low Normal Trade Relations (NTR) tariff rates the US offers to trading partners that are members of the World Trade Organization (WTO). A provision of US trade law, however, allows imports from non-market economies to enter the United States under NTR tariffs subject to *annual* approval by both the President and Congress. Chinese imports first began entering the United States under this provision in 1980 after the warming of bilateral relations. Annual approval became controversial and less certain after the Tiananmen Square incident in 1989, and this uncertainty continued throughout the 1990s. During this time, firms engaged in or considering US-China trade faced the possibility, each year, of substantial tariff increases if China's NTR status was not re-approved. The magnitude of these potential tariff increases—32 percentage points for the average product—make clear that some buyer-seller relationships that were profitable under NTR tariff rates would not be profitable under a shift to "column two" tariffs. Indeed, Pierce and Schott (2016) document extensive discussion by US firms of the trade-dampening effects of this uncertainty in the 1990s, and Handley and Limão (2017) provide a theoretical basis for these effects that operates via suppressed entry by Chinese exporters.[60] Alessandria et al. (2024) show that uncertainty regarding the annual renewal of China's NTR status each summer reduced US imports from China, while also driving intra-year seasonal patterns in imports. When the United States granted PNTR to China in 2001, it locked in NTR rates, eliminating the need for annual renewals and the potential for relationship-severing tariff increases. This plausibly exogenous policy change provides a useful opportunity for testing Proposition 2.1, i.e., whether a decrease in the probability of a trade war leads to the adoption of more "Japanese" sourcing.[61] Our strategy follows Pierce and Schott (2016) in defining a product's exposure to PNTR as the difference between

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

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

Figure S.1: Distribution of the *NTR Gap*



Source: Feenstra et al., 2002 and authors' calculations. Figure displays the distribution of the $NTR\,Gap_h$, the difference between the relatively low NTR tariff rate that was locked in by PNTR and the higher rate to which US tariffs on Chinese goods might have risen absent the change in policy.

the non-NTR rate to which its tariff could have risen before PNTR and the lower NTR rate that was locked in by the policy change,

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

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

N Supplemental DID Regressions

mhcz Quadruple Level: In the main text we show that PNTR changed the shipping patterns (quantity per shipment, weeks between shipments, and unit value) at the $mxhcz$ level. We now examine whether the shift from A to J procurement in response to PNTR also altered the shipping patterns at the $mhcz$ quadruple level. Compared to the regressions of continuing relationships at the $mxhcz$ level, this regression aggregates across the supplier dimension, and computes shipping attributes of the quadruple using transactions with all suppliers. It also allows for an additional margin of extensive margin adjustment, namely the formation of relationships with new suppliers that did not sell to the United States prior to PNTR. We use the following $mhczt$-level DID regression,

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

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

Results, displayed in Table S.11, show a significant decline in the average shipping size and weeks between shipments, consistent with a shift towards J procurement. The increase in unit values, while positive, is statistically insignificant at conventional levels. One potential explanation for this outcome is the entry of new Chinese exporters during this period (Pierce and Schott, 2016; Amiti et al., 2020), including privately owned firms that tend to have lower prices than state-owned incumbents (Khandelwal et al., 2013). New suppliers might also charge low, introductory prices to gain market share, further dampening unit values.

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

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

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

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

	(1)	(2)	(3)
Dep. var.	$\ln(QPS_{mhczt})$	$\ln(WBS_{mhczt})$	$\ln(UV_{mhczt})$
$Post_t * China_c * NTRGap_h$	-0.043***	-0.058***	0.018
	0.014	0.013	0.024
$ln(QPW_{mhczt})$	0.436***	-0.584***	-0.207***
	0.018	0.018	0.026
Observations	738,000	738,000	738,000
R-squared	0.978	0.887	0.974
Fixed effects	$mhcz, t$	$mhcz, t$	$mhcz, t$
Controls	Yes	Yes	Yes

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

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

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

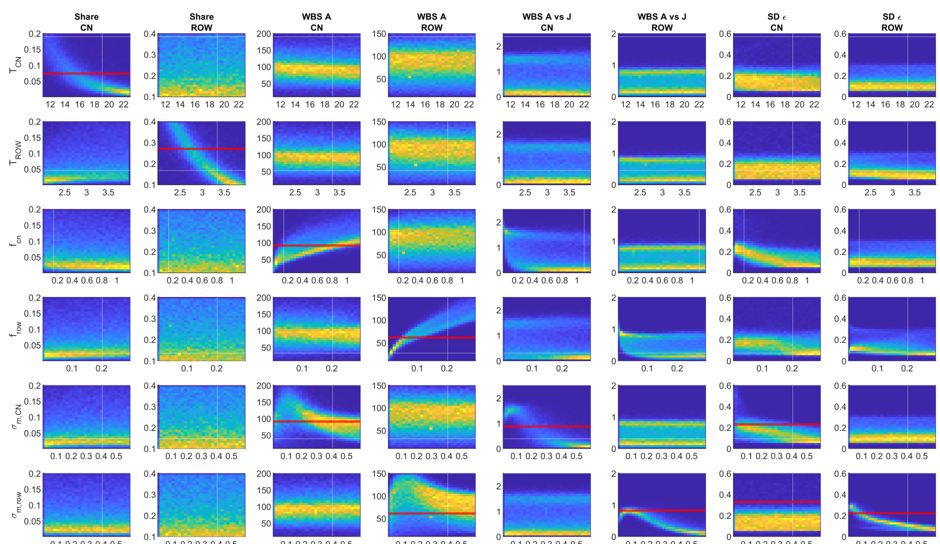
Source: LFTTD and authors' calculations. Table reports the results of regressing noted attribute of US importer by exporter by product by country by mode of transport ($mxhcz$) bins on the difference-in-differences term of interest and quantity shipped per week. Pre-and post periods are 1995 to 2000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

O Additional Quantitative Results

O.1 Identification

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

Figure S.2: Joint Identification of Parameters



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

O.2 Additional Results

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

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

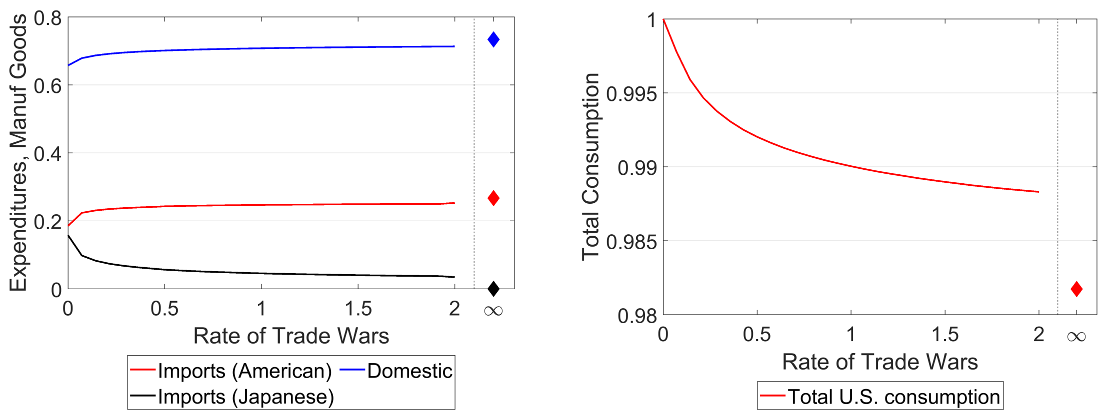
In our model the trade war arrival rate is symmetric for any country pair, and since importers always benefit from a lower arrival rate of trade wars overall welfare strictly falls with $\rho_{US,n}$, as shown in Figure S.3b. However, in a more general model in which $\rho_{n,i} \neq \rho_{i,n}$, a country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

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

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

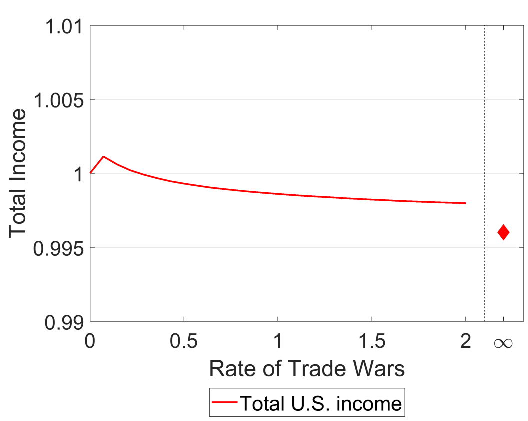
Figure S.3: Effect of Trade War Arrival Rate on Sourcing and Consumption

(a) Manufacturing Expenditures (b) Utility



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

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



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