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Trade deflection and trade depression

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Abstract

This is the first paper to empirically examine whether a country's use of an import restricting trade policy distorts a foreign country's exports to third markets. We first develop a theoretical model of worldwide trade in which the imposition of antidumping and safeguard tariffs, or "trade remedies," by one country causes significant distortions in world trade flows. We then empirically test this model by investigating the effect of the United States' use of such import restrictions on Japanese exports of roughly 4800 products into 37 countries between 1992 and 2001. Our estimation yields evidence that US restrictions both deflect and depress Japanese export flows to third countries. Imposition of a US antidumping measure against Japan deflects trade, as the average antidumping duty on Japanese exports leads to a 5–7% increase in Japanese exports of the same product to the average third country market. The imposition of a US antidumping measure against a third country depresses trade, as the average US duty imposed on a third country leads to a 5–19% decrease in Japanese exports of that same product to the average third country's market. We also document the substantial variation in trade deflection and trade depression across different importing countries and exported products.

Keywords: Deflected trade; Depressed trade; Antidumping; Safeguards; MFN

JEL classification: F10; F12; F13

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1. Introduction

In March 2002, the United States imposed a “safeguard” – a broad-based set of tariffs and quotas – on imports of steel to shield its domestic industry from foreign competition. Shortly thereafter, the European Union and a number of other steel-importing countries responded by imposing their own import restrictions on steel.[2] The EU partially justified its trade policy by arguing that the change in US trade policy would re-route or “deflect” Asian steel exports – initially destined for the newly closed US market – to what would have otherwise been a relatively open EU market.[3]

Are the EU’s concerns in the steel case consistent with historical experience? When a large importing country, such as the US, uses import restrictions such as a safeguard or antidumping duties to protect domestic producers from imports, does this lead to the substantial deflection of exports to third country markets like the EU? To our knowledge, this is the first paper to address this question empirically. We begin by presenting a simple theoretical model to illustrate the EU’s argument on deflected trade. This model embodies the potential differential impact on world trade flows of a country-specific antidumping duty (AD) versus a nondiscriminatory safeguard measure (SG).[4] We then test the model’s implications on a panel of Japanese product-level exports from 1992–2001 that is matched at the product level to changes in US trade policy through the application of antidumping duties and safeguard measures. We investigate whether there is evidence that the US use of such AD and SG “trade remedies” has an impact on Japanese export patterns to third markets and then whether there is variation across importing countries and/or products of the size of any potential distortions.

In the empirical investigation, we use a dynamic panel data model to estimate the impact of US import restrictions on Japanese exports to third countries. We construct a dataset of Japanese exports of roughly 4800 products into 37 countries between 1992 and 2001 to assess the effect of US import restrictions, thus exploiting the substantial variation across products and time of Japanese exports to third countries. Our empirical approach allows us to estimate the impact of a US-imposed, Japan-specific antidumping duty on Japanese exports, identifying whether trade is deflected to third markets. In addition we are able to identify a second impact of US antidumping duties on Japanese exports; when a US duty is applied against a third country’s exports, Japanese exports of the targeted product to the third country market are depressed.

[2] In addition to the EU, other WTO members that imposed at least preliminary safeguard protection on steel products between March 2002 and October 2003 included Chile, China, Czech Republic, Hungary, Poland and Venezuela (WTO, 2002, 2003). Canada and Bulgaria had also initiated steel safeguard investigations after the US imposition of protection but did not apply protection.

[3] The 25 March 2002 EU press release announcing its steel safeguard response to the US steel safeguard of 5 March noted that “[w]hilst US imports of steel have fallen by 33% since 1998, EU imports have risen by 18%. Given that worldwide there are 2 major steel markets (EU with 26.6 m tonnes of imports in 2001 and US with 27.6 m tonnes), this additional protection of the US steel market will inevitably result in gravitation of steel from the rest of the world to the EU. This diversion is estimated to be as much as 15 m tonnes per year (56% of current import levels).” (European Union, 2002).

[4] A ‘nondiscriminatory’ trade policy (tariff) is one that is applied equally to all exporting countries, e.g., all imports into a country face the same tariff rate. The term most-favored-nation (MFN) will be synonymous with ‘nondiscriminatory’ for the purpose of this paper. A ‘discriminatory’ or ‘preferential’ trade policy is one in which an importing country applies different tariffs to imports from different exporting countries. For example, the import tariff on goods from regional trade agreement partners is usually lower than that imposed on goods from other countries. Although the WTO requires that all its members have nondiscriminatory trade policies, there are numerous exceptions to this rule. Two of the most important exceptions in practice are that the WTO allows for discriminatory tariffs when countries participate in preferential trade agreements and when countries impose antidumping duties.

Japan is a particularly useful starting point for such an investigation for a variety of reasons. First, Japanese firms are frequently targeted by US acts of country-specific import protection; e.g., Japanese firms made up 10% of the US antidumping caseload that resulted in duties between 1992 and 2001.[5] Over the period of our sample, the US imposed antidumping duties on 157 unique 6-digit HS products from Japan. Furthermore, US import restrictions targeting third country exports affected an additional 167 products that Japan exports to one or more third countries. Second, Japanese exporters are particularly prominent in world trade. Japanese total exports as a share of world total exports was 7.5% in 1997, the midpoint of our sample.[6] Third, as Table 1 illustrates, Japanese exports to the US represent a substantial share of Japan's total exports, i.e., roughly one quarter of its total exports. This allows for the possibility of substantial trade being deflected to third country markets after the imposition of a US trade restriction.

Fig. 1 provides a preview of our empirical results on trade deflection and trade depression. The figure presents the time path of Japanese exports to third country (i.e., non-US) markets of three different categories of Japanese products: 1) those for which only Japanese firms were hit with a US antidumping measure, 2) those for which only non-Japanese firms were hit with a US antidumping measure, and 3) those for which no US antidumping measures were applied. We use the mean growth rates of these three subsamples of observations to plot the basic pattern in the data over a three year window: the year of the US antidumping investigation, as well as the two preceding years. For the Japanese products that are the target of US AD cases, there is a dramatic increase in Japanese exports to third country markets (i.e., "trade deflection") in the year of the AD investigation. Furthermore, for the Japanese exports of products to third countries where the third country exporters were the target of a US AD investigation, there is a substantial reduction (i.e., "trade depression") of Japanese exports to that third country in the year of the AD investigation. In this paper we assess whether the suggestive evidence presented in Fig. 1 is statistically and economically significant when we control for other factors affecting Japan's product-level export growth.

Our formal econometric results indicate that the imposition of US import restraints over the 1992–2001 period both deflect and depress Japanese export flows to third country markets. Imposition of a US antidumping duty against Japanese exporters is associated with substantial deflection of trade. For example, the median antidumping duty against Japan leads to a 5–7% average increase in Japanese exports to a non-US trading partner. Furthermore, there is also evidence of trade depression. When the median US antidumping duty is imposed against a third country's exporters, Japanese exports to the third country in the same product category decrease by an average of 5–19%. Finally, when faced with a US safeguard measure, Japanese exports to third countries fall by somewhere between 55% and 70%. Finally, in terms of the policy relevance, these results provide evidence that the concerns voiced by the EU in their response to the March 2002 act of US import protection may not be unfounded, given the historical experience with US trade remedies and the associated Japanese export response.

Our empirical analysis, which examines how a discriminatory trade policy change affects trade flows, fits broadly into the literature on preferential trade agreements initiated by Viner (1950). Viner identified that discriminatory trade policies associated with preferential trade agreements (PTAs) had both positive 'trade creation' welfare effects due to the enhanced trade between

[5] Japan was actually the second most targeted exporter, when measured as the number of petitions resulting in duties, as China made up 16% of the caseload.

[6] Japanese exports as a share of world total exports peaked in 1986 at 10.3%. These calculations are based on the data provided in Feenstra (2000) and include intra-EU trade in the calculation of world total exports.

Table 1
Japan's major export markets — 1996

Country	TRAINS		WTDB†	
	Rank	Export share	Rank	Export share
US	1	26.4%	1	27.5%
EU [a, b, c]	2	13.7	2	15.3
Korea [a, b, c]	3	7.7	3	7.2
Taiwan [c]	4	7.1	4	6.3
China [c]	5	7.1	6	5.5
Hong Kong [a, b, c]	6	6.4	5	6.2
Thailand [a, b, c]	7	5.7	8	4.6
Singapore [c]	8	5.4	7	5.3
Malaysia [a, b, c]	9	5.0	9	3.9
Canada [a, b, c]	10	2.5	13	1.4
Indonesia [a, b, c]	11	2.2	10	1.9
Philippines [c]	12	2.0	12	1.7
Australia [a, b, c]	13	2.0	11	1.8
Brazil [a, b, c]	14	0.8	17	0.6
Mexico [a, b, c]	15	0.8	15	0.7
New Zealand [a, b, c]	16	0.6	22	0.4
Saudi Arabia [c]	17	0.6	16	0.7
Switzerland [a, b, c]	18	0.5	19	0.5
India [a, b, c]	19	0.5	20	0.5
Turkey [a, b, c]	20	0.4	24	0.3
Norway [a, b, c]	21	0.3	28	0.3
South Africa [a, b, c]	22	0.3	21	0.5
Israel [c]	23	0.3	26	0.3
Russia [a, b, c]	24	0.3	23	0.3
Chile [c]	25	0.2	29	0.2
Argentina [a, b, c]	26	0.2	36	0.1
Colombia [a, c]	27	0.2	32	0.2
Peru [a, c]	28	0.1	41	<0.1
Bangladesh [a, b, c]	29	0.1	42	<0.1
Poland [a, b, c]	30	0.1	40	<0.1
Czech Republic [a, b, c]	31	0.1	39	<0.1
Egypt [a, b, c]	32	<0.1	37	0.1
Hungary [a, b, c]	33	<0.1	45	<0.1
Venezuela [a, c]	34	<0.1	44	<0.1
Nigeria [c]	35	<0.1	43	<0.1
Panama [a, c]	36	<0.1	14	1.3
Guatemala [c]	37	<0.1	76	<0.1
Romania [a, c]	38	<0.1	85	<0.1

Notes: † World Trade Data Base source is Feenstra (2000).
[a] Japanese trading partner used in estimating specifications (1) and (6).
[b] Japanese trading partner used in estimating specifications (2), (3), (4), and (5).
[c] Japanese trading partner used in estimating specifications (7), (8), (9), and (10).

members (allowing members to exploit comparative advantage amongst themselves) and negative 'trade diversion' welfare effects by potentially reducing trade between members and non-members (and thus preventing the full exploitation of worldwide comparative advantage). Ultimately, Viner recognized that the overall welfare effect of a PTA would have to be assessed empirically. More recently, a substantial theoretical literature (including, but not limited to Bond and Syropolous, 1996; Bagwell and Staiger, 1997, 1999, 2004; Levy, 1997; Ethier, 2004;

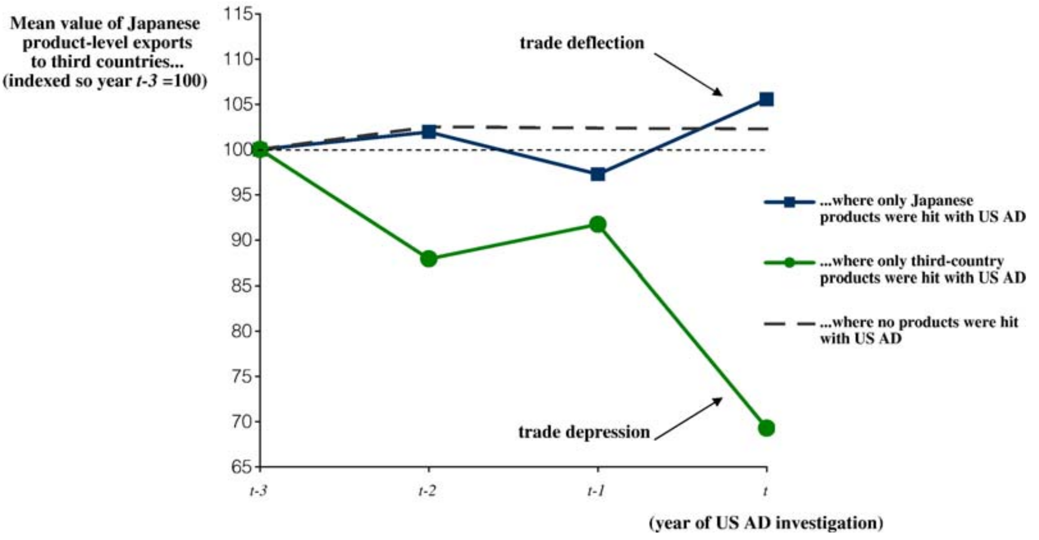

Fig. 1. Trade deflection and trade depression associated with US antidumping.

McLaren, 2002) examines the role of preferential policy exceptions in multilateral trade agreements.[7] These papers typically focus on the *import source* diversion as the mechanism through which discriminatory trade policies affect welfare; the domestic welfare losses are derived from importing from someone who is not the lowest cost producer and failing to (globally) exploit comparative advantage. In contrast, our analysis will focus on *export* diversion where global welfare costs would arise because the low cost exporter is being shut out of a market for which it would potentially be the most efficient producer if there were non-discriminatory application of tariffs.

Our empirical approach is most similar to Romalis (2002) which investigates the import source diversion of Mexican and Canadian exports to the US resulting from the North American Free Trade Agreement (NAFTA) and the earlier Canada–US Free Trade Agreement (CUSFTA), respectively.[8] Romalis uses a similarly disaggregated panel of product-level export data and finds that Mexican and Canadian shares of US imports have increased most rapidly in the products facing the largest changes in trade policy; i.e., where the greatest PTA tariff preferences were conferred. While not the focus of his analysis, in presenting information on two of his controls, Mexican and Canadian shares of EU imports, he documents evidence that is consistent with our results regarding the deflection of trade.[9] Although related, our paper differs from Romalis' both

[7] For a survey of other recent papers focusing on different theoretical elements of the interaction between preferential and multilateral agreements, see Krishna (2004). For a literature survey on the nondiscrimination principle and the economic aspects of the MFN clause in trade agreements, see Horn and Mavroidis (2001).

[8] Clausing (2001) is another recent paper that looks at the trade creation and trade diversion effects of the CUSFTA through an analysis of a panel on product-level trade data and tariff changes. Romalis (2002) presents a thorough discussion of the differences in approaches and results of the two papers.

[9] We interpret Romalis' Fig. 1B and C for Mexico and Fig. 2B and C for Canada as evidence of the deflection of exports from the EU to the US. Starting from a non-discriminatory benchmark, the discriminatory removal of US trade barriers lowers tariffs facing a Mexican or Canadian product, and leads to trade being deflected away from a third market like the EU.

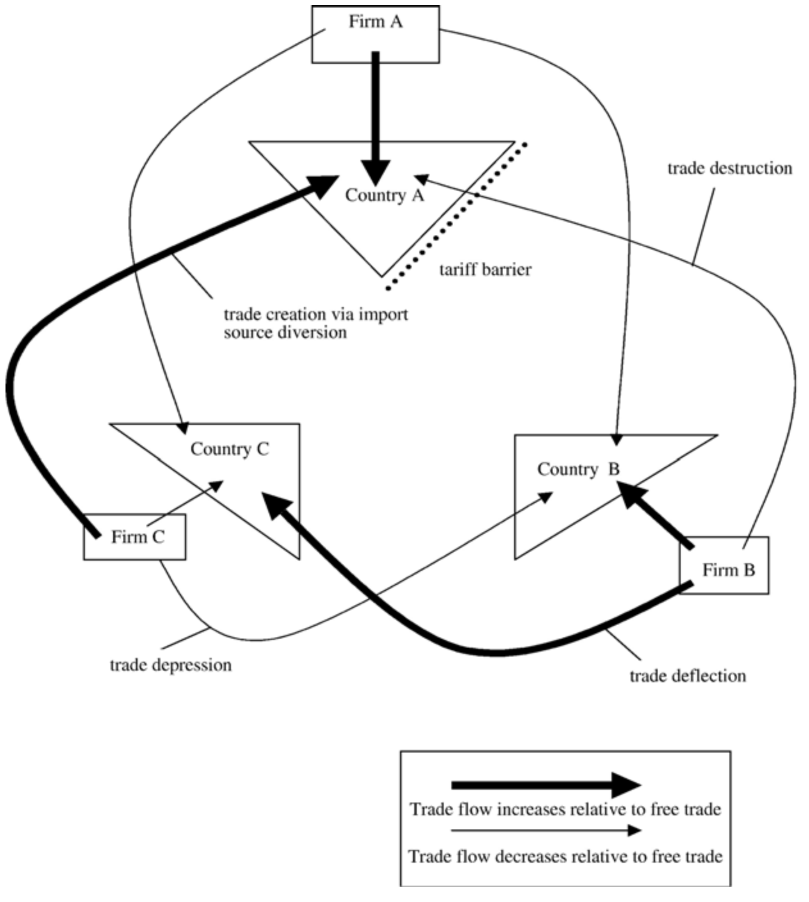

Fig. 2. Trade flows under an antidumping duty.

in terms of the question we ask and the empirical methodology we employ. Romalis uses a difference-in-differences approach to examine if changes in US trade policy (NAFTA and CUSFTA) produced import source diversion. We use dynamic panel data methods to examine if changes in US trade policy (AD and SG) are associated with the *deflection* and also the *depression* of exports.

With respect to the economics literature on trade remedies, our results on the trade distorting effects of antidumping measures complement the work of Prusa (1997, 2001).[10] Prusa (2001), for example, uses a panel of US industry-level imports and data on US antidumping measures for 1980–1994 to investigate the import source diversion that occurs for the United States when a discriminatory trade policy causes importers to switch from a lower cost to a higher cost foreign supplier. He provides evidence that foreign exporters in an industry subject to a US AD measure who are not-named in an antidumping petition increase exports to the US in conjunction with the exports of the country targeted by the AD petition falling. Our paper can provide insight as to

[10] For a recent survey of the economics literature on antidumping, see Blonigen and Prusa (2004). For a survey on safeguards protection, see Bown and Crowley (2005).

where the products targeted by the US petition go, since they are no longer being exported to the US market.[11]

The rest of this paper proceeds as follows. Section 2 presents our simple economic model to flush out our empirical predictions. Section 3 presents our empirical model that will be used in the estimation and a discussion of variable construction and data. In Section 4 we discuss our empirical results, and Section 5 concludes with a discussion of additional questions and further puzzles for future research.

2. Theoretical model

Assume there are three countries indexed i or $j \in \{A, B, C\}$, $i \neq j$. Each country has one firm, also indexed i or j, which produces a single good for domestic consumption and for export. A good is denoted m_{ij}, where the first index, i, indicates the country of production, and the second index, j, indicates the country in which the good is consumed. Thus, a good produced by firm i for export to country j is denoted m_{ij}. Output produced for domestic consumption is denoted m_{ii}. Markets are segmented, firms compete on quantity, and the good produced for domestic consumption and the imported goods are strategic substitutes ($\pi_{m_{ii}m_{ji}}<0$, $\pi_{m_{ji}m_{-ji}}<0$).

Production in each country employs the same technology. The marginal cost of production is increasing, the cost function is $c(x_i)$ where $c'(x_i)>0$ and $c''(x_i)>0$ and x_i is firm i's total output. Firm i's total output is the sum of domestic sales and sales in the two foreign markets, $x_i = \Sigma_j m_{ij}$, $j \in \{A, B, C\}$.

Inverse demand in all countries is given by $p(Q_i, Y_i)$ where Q_i is the total output sold in country i and Y_i is national income. Total output sold in i is the sum of domestic sales by the domestic firm and imports from the other two countries, $Q_i = \Sigma_j m_{ji}$, $j \in \{A, B, C\}$.

The objective of the firm in i is to chose a total output level and a level of sales for each market in order to maximize profits,

$$\max_{m_{ij}} \pi_i = \sum_j [p(Q_j)m_{ij} - \tau_{ij}m_{ij}] - c(x_i), \tag{1}$$

where τ_{ij} represents country j's tariff on imports from i and τ_{ii}, the tariff on consumption of the domestically produced good, is equal to zero. The firms' first order conditions are given by the following:

$$\frac{\partial \pi_i}{\partial m_{ij}} = p(Q_j) + p'(Q_j)m_{ij} - \tau_{ij} - c'(x_i) = 0. \tag{2}$$

Solving the first order conditions for each $j \in \{A, B, C\}$ yields firm i's best responses to the sales decisions of the other two firms. A best response function specifies an amount to sell in each market, given the sales in that market of the firm's two rivals. Solving the nine best response functions simultaneously yields the Cournot Nash equilibrium quantities sold by each firm in each country.

$$m_{ij} = f(p(Q_i, Y_i), c(x_i), \tau_{ij}) \quad \forall i, j \in \{A, B, C\} \tag{3}$$

[11] While we do not investigate the issue here, our results suggest that there may also be substantial welfare distorting effects of acts of US trade protection outside of the US, in addition to the sizable welfare distortions experienced inside the US and documented, for example, by Gallaway, Blonigen and Flynn (1999).

In the Cournot Nash equilibrium, because the marginal cost of production is increasing, each firm will choose to allocate its total output across the three countries so that its net marginal revenue (marginal revenue less tariff costs) is the same in all three markets.

2.1. Comparative statics for an antidumping duty

Without loss of generality, suppose that trade among the three countries is free, with the exception that country A imposes a tariff on imports from country B. How will an increase in A's tariff affect trade among all three countries? Fig. 1 provides an illustration of Proposition 1.

Proposition 1. *For the three country Cournot model in which goods are strategic substitutes and firms face increasing marginal costs in production, a tariff by country A against country B causes, relative to the free trade equilibrium:*

1. *trade destruction, a decline in country B's exports to country A* $\left(\frac{\mathrm{d}m_{\mathrm{ba}}}{\mathrm{d}\tau_{\mathrm{ba}}}<0\right)$,
2. *trade creation via import source diversion, an increase in country C's exports to country A* $\left(\frac{\mathrm{d}m_{\mathrm{ca}}}{\mathrm{d}\tau_{\mathrm{ba}}}>0\right)$,
3. *trade deflection, an increase in country B's exports to country C* $\left(\frac{\mathrm{d}m_{\mathrm{bc}}}{\mathrm{d}\tau_{\mathrm{ba}}}>0\right)$, *and*
4. *trade depression, a decrease in country C's exports to country B* $\left(\frac{\mathrm{d}m_{\mathrm{cb}}}{\mathrm{d}\tau_{\mathrm{ba}}}\right)<0$.

Proof. Totally differentiating the nine first order conditions given by (2), dividing through by $\mathrm{d}\tau_{ab}$, and applying Cramer's rule yields the signs of the comparative static effects on the domestic output and exports of all three firms of an increase in country A's tariff on imports from country B. For strategic substitutes and an increasing marginal cost of production, without loss of generality, the following results are obtained for a change in τ_{ba}: for goods consumed in country A, $\frac{\mathrm{d}m_{\mathrm{aa}}}{\mathrm{d}\tau_{\mathrm{ba}}}>0$, $\frac{\mathrm{d}m_{\mathrm{ba}}}{\mathrm{d}\tau_{\mathrm{ba}}}<0$, $\frac{\mathrm{d}m_{\mathrm{ca}}}{\mathrm{d}\tau_{\mathrm{ba}}}>0$, for goods consumed in country B, $\frac{\mathrm{d}m_{\mathrm{bb}}}{\mathrm{d}\tau_{\mathrm{ba}}}>0$, $\frac{\mathrm{d}m_{\mathrm{ab}}}{\mathrm{d}\tau_{\mathrm{ba}}}<0$, $\frac{\mathrm{d}m_{\mathrm{cb}}}{\mathrm{d}\tau_{\mathrm{ba}}}<0$, for goods consumed in country C, $\frac{\mathrm{d}m_{\mathrm{cc}}}{\mathrm{d}\tau_{\mathrm{ba}}}<0$, $\frac{\mathrm{d}m_{\mathrm{bc}}}{\mathrm{d}\tau_{\mathrm{ba}}}>0$, $\frac{\mathrm{d}m_{\mathrm{ac}}}{\mathrm{d}\tau_{\mathrm{ba}}}<0$. □

In this model, the existence of a deflected trade flow relies critically on the assumption of an increasing marginal cost of production. Because firms equate the net marginal revenue of producing for each market in equilibrium, anything that raises the cost of selling in one market will cause firms to reallocate their sales across markets.

2.2. Comparative statics for a safeguard tariff

Without loss of generality, suppose that trade among the three countries is free, with the exception that country A imposes a tariff on imports from countries B and C. Assume that the magnitudes of the tariffs set against B and C are identical ($\tau_{\mathrm{ba}}=\tau_{\mathrm{ca}}$) and given by τ. How will an increase in country A's tariff affect trade among all three countries?

Proposition 2. *For the three country Cournot model in which goods are strategic substitutes and firms face increasing marginal costs in production, a tariff by country A against all other countries (B and C) causes, relative to the free trade equilibrium:*

1. *trade destruction, a decline in country C and B's exports to country A* $\left(\frac{\mathrm{d}m_{\mathrm{ba}}}{\mathrm{d}\tau}<0, \frac{\mathrm{d}m_{\mathrm{ca}}}{\mathrm{d}\tau}<0\right)$ *and*
2. *two-way trade deflection, an increase in country B's exports to country C* $\left(\frac{\mathrm{d}m_{\mathrm{bc}}}{\mathrm{d}\tau}>0\right)$ *and an increase in country C's exports to country B* $\left(\frac{\mathrm{d}m_{\mathrm{cb}}}{\mathrm{d}\tau}>0\right)$.

Proof. Totally differentiating the nine first order conditions given by (2), dividing through by $\mathrm{d}\tau$, and applying Cramer's rule yields the signs of the comparative static effects of an increase in country A's tariff on imports from all countries on the domestic output and exports of all three firms. For strategic substitutes and an increasing marginal cost of production, without loss of generality, the following results are obtained for a change in τ: for goods consumed in country A, $\frac{\mathrm{d}m_{\mathrm{aa}}}{\mathrm{d}\tau}>0$, $\frac{\mathrm{d}m_{\mathrm{ba}}}{\mathrm{d}\tau}<0$, $\frac{\mathrm{d}m_{\mathrm{ca}}}{\mathrm{d}\tau}<0$, for goods consumed in country B, $\frac{\mathrm{d}m_{\mathrm{bb}}}{\mathrm{d}\tau}>0$, $\frac{\mathrm{d}m_{\mathrm{ab}}}{\mathrm{d}\tau}<0$, $\frac{\mathrm{d}m_{\mathrm{cb}}}{\mathrm{d}\tau}>0$, for goods consumed in country C, $\frac{\mathrm{d}m_{\mathrm{cc}}}{\mathrm{d}\tau}>0$, $\frac{\mathrm{d}m_{\mathrm{bc}}}{\mathrm{d}\tau}>0$, $\frac{\mathrm{d}m_{\mathrm{ac}}}{\mathrm{d}\tau}<0$. □

Comparing a discriminatory antidumping policy and a nondiscriminatory safeguard, the theoretical model predicts that two phenomena observed under an antidumping duty – trade creation via import source diversion and trade depression – are absent under a safeguard. Because a safeguard creates an identical increase in costs on products from both import sources, there is no incentive to favor one source over another. Thus, the result that no trade is created through import source diversion is fairly obvious. With regard to the model's prediction of two-way trade deflection under a safeguard, the result is less obvious. For each country B and C, the safeguard induces two conflicting forces of trade depression and trade deflection. Retained domestic production that can no longer be sold in country A could "crowd out" imports and lead to trade depression, but in the model this effect is swamped by each firm's strong desire to export so that it will not be competing against itself in its domestic market.

In the next section, we test the model's predictions about trade deflection, trade depression, and two-way trade deflection on a panel of Japanese product exports. Our approach is thus different from papers by Romalis (2002) and Prusa (1997, 2001) who estimate empirical models of what we refer to as "trade destruction" and "trade creation via import source diversion,"[12] respectively.

3. Empirical model and estimation

3.1. The empirical investigation

The theoretical model presented in Section 2 yields a number of predictions relating one country's tariffs to trade flows between foreign countries. Our empirical analysis focuses on the predictions of deflected, depressed and two-way deflected trade for Japanese exports to 37 non-US trading partners. For clarity of exposition, ignoring Japan's 36 other trading partners, what does our theoretical model predict when the country imposing tariffs is the US and the foreign countries are Japan and the EU? First, if the US imposes a country-specific tariff against Japan in the form of an antidumping duty and imposes no tariff against the EU, the model predicts deflected trade, an increase in Japanese exports to the EU. Second, if the US imposes a country-specific tariff against the EU in the form of an antidumping duty, but not against Japan, the model predicts that Japanese exports to the EU will fall, i.e., depressed trade. In this case, European exports that are diverted away from the US market by the tariff and sold domestically within the EU depress imports from Japan. Third, if the US imposes tariffs against both Japan and the EU in the form of a broadly-applied safeguard measure or two simultaneously-imposed antidumping duties, the model predicts two-way deflected trade, a rise in Japanese exports to the EU and in EU exports to Japan.[13]

[12] More precisely, Romalis estimates the opposite effect - trade creation arising from the removal of a tariff.

[13] We do not test for the rise in EU exports to Japan here as our analysis focuses on the response of Japanese exports only.

3.2. Basic empirical model

To investigate the questions identified by the theoretical model, we develop the following reduced-form specification for the value of Japanese exports to country i based on Eq. (3):[14]

$$\begin{aligned}\ln(vm_{iht}) = {} & \alpha_i + \gamma_h + \beta_1'\ln(Y_t) + \beta_2'\ln(Y_{it}) + \beta_3'\ln(e_{it}) + \beta_4'\tau_{ht} + \beta_5'\tau_{iht} + \beta_6'\tau_t^{*i} \\ & + \beta_7'\ln(c_{kt}) + \beta_8'\ln(vm_{iht-1}) + \epsilon_{iht},\end{aligned} \tag{4}$$

where i denotes an importing country, h denotes a 6-digit HS product, and t denotes time in years. The index k denotes an industry aggregate at the 3-digit ISIC level, i.e. the products $h=1\ldots h'$ map into the industries $k=1$, $h=h'\ldots h''$ map into $k=2$, and so on until $h=h^*\ldots H$ map into $k=K$.

The variable vm_{iht} denotes the value of imports from Japan of h into i at time t, Y_t denotes Japan's national income (an export-supply shifter), Y_{it} denotes the importing country i's national income (an import-demand shifter) and e_{it} is the exchange rate between the yen and the importing country's currency. The variable τ_{ht} designates US trade policy against Japan while τ_{iht} captures US trade policy against importing country i. Japan's industry k cost variables are denoted by c_{kt}, while importing country i's trade policy is denoted by τ_t^{*i}.[15] Finally, α_i denotes country fixed effects, γ_h denotes product fixed effects, the β's are the parameters to be estimated, and ϵ_{iht} is the error term.

3.3. Estimation strategy

There are two problems to address in estimating Eq. (4). First, the autocorrelation of vm_{iht} implies that least squares estimation of Eq. (4) yields biased estimates. Second, in a short panel, the number of parameters to be estimated (α_i and γ_h) increases with the number of countries and products. Thus, α_i and γ_h cannot be consistently estimated.

To address both of these problems, we estimate the first difference of Eq. (4) using the optimal Generalized Method of Moments (GMM) estimator proposed by Arellano and Bond (1991), in which multiple lags of the level of the dependent variable are used as instruments for lags of the first difference of the dependent variable.[16] We thus use GMM to estimate

$$\begin{aligned}\Delta\ln(vm_{iht}) = {} & \beta_1'\Delta\ln(Y_t) + \beta_2'\Delta\ln(Y_{it}) + \beta_3'\Delta\ln(e_{it}) + \beta_4'\Delta\tau_{ht} + \beta_5'\Delta\tau_{iht} + \beta_6'\Delta\tau_t^{*i} \\ & + \beta_7'\Delta\ln(c_{kt}) + \beta_8'\Delta\ln(vm_{iht-1}) + \Delta\epsilon_{iht}.\end{aligned} \tag{5}$$

3.4. Fixed effects model

One potential criticism of our basic empirical model is that it does not adequately control for product-level variation in production costs because our industry cost variables, c_{kt} , are only available at a 3-digit industry level whereas our trade data and policy changes are measured at a 6-digit product level. Therefore, as a robustness check, we first difference Eq. (4) and use 6-digit HS product fixed effects and country–year dummies to control for detailed product-level

[14] Unfortunately, only the value of imports is consistently available in the TRAINS data, so we cannot analyze the price and quantity responses to a trade policy change separately.

[15] Ultimately it would be preferable to also have product-level data for country i's trade policy and Japan's costs; unfortunately neither of which is yet systematically available over this time period.

[16] Direct estimation of the first difference of Eq. (4) by least squares would yield biased coefficients because the lagged difference of imports $[\ln(vm_{iht-1})-\ln(vm_{iht-2})]$ is correlated with the error term $[\epsilon_{iht}-\epsilon_{iht-1}]$.

variation as well as country-specific macroeconomic variation over time. Our fixed effects model is given by the following:

$$\Delta\ln(vm_{iht}) = \mu_h + \chi_{it} + \eta_1'\Delta\tau_{ht} + \eta_2'\Delta\tau_{iht} + \eta_3'\Delta\ln(vm_{iht-1}) + \Delta\epsilon_{iht}, \tag{6}$$

where μ_h are 6-digit HS product-specific fixed effects and χ_{it} represents a full set of country–year dummies. Because of the large number of parameters to be estimated, Arellano and Bond's GMM estimator is not computationally feasible. Therefore, we estimate Eq. (6) using the two stage least squares/instrumental variables (IV) approach of Anderson and Hsiao (1981, 1982) in which we instrument for $\Delta\ln(vm_{iht-1})$ using the second lag of the log level of the value of imports. Because the product fixed effects and the country–year dummies absorb product-level and macroeconomic variation over time, this approach requires fewer control variables (e.g. GDP growth, value-added per worker) than estimation of Eq. (5) and thus we are able to utilize a much larger sample of trade and trade remedy data for many additional countries.

Nevertheless, because of the dynamic panel structure of our data, there are two potential problems with the Anderson and Hsiao (1981, 1982) IV estimator; bias associated with the use of a weak instrument and bias associated with correlation in measurement error. In Appendix A we address both of these concerns. To address the weak instrument problem, we test the quality of two instruments, $\ln(vm_{iht-2})$ and $\ln(vm_{iht-3})$. We find that both are strong instruments for $\Delta\ln(vm_{iht-1})$ and conclude the IV approach is appropriate for our problem. To address the issue of measurement error, we compare coefficient estimates using the second and third lags of the log level of imports. We find that our coefficient estimates are robust to the choice of instrument, suggesting that measurement error is not a significant problem and the use of $\ln(vm_{iht-2})$ as an instrument is appropriate.

3.5. Variable construction and data

In this section we discuss the construction of variables used in the estimation of Eqs. (5) and (6) as well as the sources of our data. Table 2 summarizes variable descriptions and our predictions about the signs of the estimated coefficients, as well as providing summary statistics.

3.5.1. Trade variables

First consider the dependent variable in the estimation of Eqs. (5) and (6), $\Delta\ln(vm_{iht})$, which is the annual growth of third country i's imports of product h from Japan. The detailed, highly disaggregated data used in this paper represent a significant improvement over many previous studies on US import restrictions and trade remedies. Annual data on the nominal value of imports into 37 non-US countries for roughly 4800 6-digit Harmonized System (HS) products for the years 1992 to 2001 come from UNCTAD's TRAINS data base. Import data for these 37 countries was reformatted into a dataset of Japanese exports to these countries.[17] In our basic specification of Eq. (5), we are restricted to using a smaller set of 28 importing third countries due to the limited

[17] Because this data is collected only on the import side, it is possible that discrepancies exist between a country's imports from Japan and Japan's exports to that country. We checked the quality of our Japanese export dataset against Feenstra's (2000) NBER's World Trade Database (WTDB). The WTDB includes data on worldwide import and export flows at a 4-digit SITC level and is thus too aggregated for our purposes, but is known to be of high quality because it matches import and export records to resolve any discrepancies in the values of trade flows between pairs of countries. Table 1 presents a comparison of Japan's aggregate export shares in 1996 calculated using our dataset and the WTDB. The shares from the two datasets are comparable and we feel confident in using the TRAINS data in our analysis. Nevertheless there are some years for which trade data is missing for certain countries.

Table 2
Data description and summary statistics

Variable	Vector	Description	Predicted sign	Mean	Standard deviation	Observations
Dependent variable						
$\Delta\ln(vm_{iht})$	$\Delta\ln(vm_{iht})$	Growth of country i's real imports of product h from Japan†		−0.0250	0.9532	254,074
Explanatory variables						
ΔAD $\text{duty}_{\text{jpn},ht}$	$\Delta\tau_{ht}$	US AD duty against Japan on product h	(+)	0.0014	0.0324	254,074
ΔAD duty_{iht}	$\Delta\tau_{iht}$	US AD duty against country i on product h	(−)	0.0003	0.0187	254,074
ΔSG policy_{ht}	$\Delta\tau_{ht}$	US SG policy on product h in year t	(+)	0.0001	0.0095	254,074
$\Delta\ln(\text{realGDP})_{it}$	$\Delta\ln(Y_{it})$	Growth of country i's GDP	(+)	0.0273	0.0451	141,164
$\Delta\ln(\text{realGDP})_{\text{jpn},t}$	$\Delta\ln(Y_t)$	Growth of Japan's GDP	(−)	0.0111	0.0148	141,164
$\Delta\ln(\text{open})_{it}$	$\Delta\tau_t^{*i}$	Growth of country i's openness to world $[\Delta\ln((X_{it}+M_{it})/\text{GDP}_{it})]$	(+)	0.0311	0.0849	141,164
$\Delta\ln(\text{yen/curr}_i)_t$	$\Delta\ln(e_{it})$	Growth of bilateral real Japanese yen/country i's currency rate*	(+)	0.0011	0.1465	136,583
$\Delta\ln(\text{avg.wage})_{\text{jpn},kt}$	$\Delta\ln(c_{kt})$	Growth of Japan's industry k average wage	(+)	−0.0297	0.1182	113,393
$\Delta\ln(\text{v.add/worker})_{\text{jpn},kt}$	$\Delta\ln(c_{kt})$	Growth of Japan's industry k average value-added per worker	(+)	−0.0247	0.1419	113,393

Notes: Subscript i is an importing country, h is a 6-digit HS product, k is a 3-digit ISIC industry, and t is a year. † The trade-weighted mean growth of country i's real, product-level imports from Japan is 0.1743 with a trade-weighted standard deviation of 0.6222 in the sample of 254,074 observations, and 0.1684 with a trade-weighted standard deviation of 0.6174 in the sample of 141,164 observations. *For Norway and Switzerland, this is the growth of the real exchange rate index from the IFS.

availability of some of the macroeconomic data needed for the estimation. The alternative fixed effects model Eq. (6) requires no macroeconomic data and utilizes a larger sample of 37 importing countries. The countries in the final dataset include OECD members, many countries from Asia and Latin America, and some former members of the USSR and Eastern European countries. Data on Africa is generally not available in TRAINS, but as these countries are extremely small markets for Japanese exports, their omission should not affect our results. Table 1 also lists the countries used to estimate the different specifications. Because the TRAINS dataset does not include product-specific price deflators, we deflated the nominal import data, which is reported in US dollars, using the US Bureau of Labor Statistic's HS Import Price Indices, which are available for the period 1992 through 2001.

3.5.2. US antidumping and safeguard policy variables

The main explanatory variables of interest in Eqs. (5) and (6) are the changes to US import policy facing a product h exported to the US from Japan ($\Delta\tau_{ht}$) or from a third country ($\Delta\tau_{iht}$). Our estimates use data on the country-specific, trade-weighted average of the antidumping duty in the year in which the antidumping measure was imposed.[18] For a product h, we examine the effect of (1) the imposition of a US antidumping duty against Japan, (2) the imposition of a US antidumping duty against a third country i, and (3) the imposition of a US safeguard policy. As

[18] The duty data was generously provided by Bruce Blonigen and his AD website http://darkwing.uoregon.edu/bruceb/adpage.html.

discussed in Section 2, the theoretical model predicts that the sign of the coefficient on (1) is positive, on (2) is negative, and on (3) is positive.[19]

We collected data on the US imposition of country-specific antidumping duties and safeguard measures at the 6-digit HS level over the 1992 through 2001 period from a variety of US government publications, most notably the Federal Register. For antidumping and safeguard cases filed during the sample, we obtained the names and 6-digit HS codes for the products involved, the outcome of the case (affirmative, negative, or terminated, as well as the type of measure for safeguards), the names of the countries that faced the import restrictions, the trade-weighted average duty when duties were imposed, and, most importantly, the date a case was initiated and the date a trade restriction began.[20] For the antidumping policies, we interact a variable indicating that the policy was imposed in year t with the level of the antidumping duty that is imposed, to help control for the heterogeneity in duties imposed across exporters and across investigations. On the other hand, we use a simple indicator variable to examine the safeguard policies, due to the fact that sometimes the safeguard measure is imposed as a quantitative restriction or a tariff-rate quota, as opposed to a simple ad valorem duty.

3.5.3. Macroeconomic variables

We include controls for growth in the exporting ($\Delta\ln(Y_t)$) and importing countries' GDP ($\Delta\ln(Y_{it})$), growth of the real yen/country i currency exchange rate ($\Delta\ln(e_{it})$), and proxies for changes in the importing countries' trade policies ($\Delta\tau_t^{*i}$).

We expect an increase in the GDP growth of the exporting country (Japan) to lead to a fall in Japanese export growth because domestic demand for the export goods will be higher. In other words, Japan is expected to export domestic weakness. Second, in terms of currency changes, export growth should be higher when the yen is weakening relative to the importing country's currency. Thus, we expect a positive sign on the coefficient for growth of the real exchange rate.

For the importing country, an increase in GDP growth should be associated with higher Japanese export growth. To proxy for changes to an importing country's overall trade policy that we cannot observe, for example, an across the board tariff reduction or a reduction in the administrative cost of exporting to a particular importing country, we control for changes in an importing country's "openness." Openness is defined as the sum of real aggregate imports and exports divided by real GDP. For some countries, real aggregate import and export series were not available. For these countries, we calculated "openness" using the corresponding nominal

[19] We do not investigate the impact on Japanese exports of US AD (or SG) investigations that do not result in duties, but which are terminated or settled. While such an investigation could lend further insight into the overall impact of Staiger and Wolak's (1994) "investigation effect," "suspension effect" and "withdrawal effect," of the non-duty impact of AD investigations, it is beyond the scope of issues under investigation here and thus we leave it to future research. Furthermore, while the theoretical model also generates predictions on the expected impact of the removal of trade remedies, we do not report estimates of those impacts here. The primary issue is the quality of the antidumping policy removal data. We are primarily concerned with measurement error as it is difficult to "time" when an antidumping measure is removed, given that in many cases the removal is applied retroactively (with the refunding of duties), but in a manner which could not have been reasonably anticipated by the exporters. Nevertheless, we do note that the estimates (available from the authors upon request) for our policy variables of interest are not significantly altered with the inclusion of the policy removal variables.

[20] To clarify the timing of our different variables, the variables $\Delta\text{AD Duty}_{jpn,ht}$ and $\Delta\text{AD Duty}_{iht}$ are nonzero in the period in which the investigation into an antidumping case that results in a duty is begun. The $\Delta\text{SG Policy}_{ht}$ variable is an indicator that is equal to 1 in the period in which a safeguard measure goes into effect. This reflects the fact that in AD cases the targeted exporters begin to respond to provisional antidumping duties that are imposed shortly after the date the investigation is announced. Safeguard cases, on the other hand, have a very uncertain outcome and almost never use temporary trade restrictions during the investigation phase.

variables. We believe that an increase in this variable is associated with liberalization of country i's trade policy, and thus, expect a positive sign on its coefficient.

Data on real GDP, real aggregate imports, and real aggregate exports come from two sources: the OECD Main Economic Indicators and the IMF's International Financial Statistics (IFS). Whenever possible, we used the OECD data to construct the macroeconomic controls. When OECD data were not available or were only available for a short timespan, we used data from the IFS. We construct real bilateral Japanese Yen to country i currency rates for 20 countries using data supplied by the USDA Economic Research Service. An increase in the value of the real exchange rate implies an appreciation of country i's currency. For Norway and Switzerland, bilateral rates were not available from USDA so we use real exchange rate indices from the IFS.[21]

3.5.4. Industry-level variables

Lastly, we use two measures of productivity changes for Japanese manufacturing industries: the growth of the average wage and the growth of value-added per worker. This addresses a concern that our policy variables may not be measuring true treatment effects, but may be picking up the effect of an omitted variable – like a Japanese productivity improvement – that would be associated with the imposition of a US import barrier on Japanese imports and an increase in Japanese export growth to other countries. We expect the sign on both productivity measures to be positive.

Japanese manufacturing industry data at the 3-digit ISIC (Rev. 2) level for the years 1992–1999 came from the UNIDO (2002). We used data on number of employees, value-added and average wages to construct two productivity measures: the growth of value-added per worker and the growth of average wages.

4. Empirical results

4.1. Estimation results using the GMM procedure

Table 3 presents our estimates of Eq. (5) using the Arellano and Bond (1991) GMM estimator. Specifications (1) and (2) present estimates on the full set of industries (agricultural and manufacturing) over the 1992–2001 period thus leaving out the industry-level controls. Specifications (3) through (5) present estimates for all manufacturing industries from 1992–1999, all years for which the ISIC industry variables are available.

Consider first specification (1) and our estimates for the policy variables of interest, which provide evidence in support of some of the key predictions of our theoretical model. US imposition of antidumping measures against Japan is associated with statistically significant *deflection* of Japanese exports to third country markets, and US imposition of antidumping measures against third countries is associated with a statistically significant *depression* of Japanese exports to those markets. Interestingly, the coefficient estimate on the safeguard indicator variable also implies that a safeguard policy is associated with a statistically significant depression of Japanese exports to third markets.

With respect to the size of the estimates, specification (1) indicates that the imposition of a 1% antidumping duty against Japanese exports of product h (but not exporters from country i) is associated with an 0.14% *increase* in Japanese exports of h to country i. To understand the magnitude of the effect, consider that the median antidumping duty (conditional on a duty being

[21] For the IFS series, an increase in the value of the real exchange rate index implies a real appreciation of the Norwegian and Swiss currencies, respectively. We thank Matthew Shane of the ERS of the USDA for providing us with the data and answering questions about the construction of the USDA's real exchange rates.

Table 3
Arellano–Bond GMM estimation: results for Japanese exports, 1995–2001

Explanatory variables	Dependent variable: $\Delta\ln(vm_{iht})$					
	Baseline specification (1)	Add currency depreciation (2)	Add industry controls (3)	Substitute growth in value-added per worker (4)	Remove industry controls (5)	IV/fixed effects specification instead of GMM (6)
Policy variables						
ΔAD duty$_{\text{jpn},ht}$	0.141* (0.086)	0.148* (0.086)	0.156* (0.091)	0.168* (0.091)	0.167* (0.090)	0.127 (0.081)
ΔAD duty$_{iht}$	−1.269*** (0.411)	−1.207*** (0.408)	−1.292*** (0.418)	−1.262*** (0.417)	−1.265*** (0.415)	−0.870*** (0.243)
ΔSG policy$_{ht}$	−1.012* (0.536)	−0.981* (0.530)	−1.023 (2.111)	−1.102 (2.105)	−1.018 (2.099)	−1.189*** (0.275)
Other control variables						
IV for $\Delta\ln(vm)_{iht-1}$	0.276*** (0.012)	0.245*** (0.012)	0.328*** (0.012)	0.319*** (0.012)	0.306*** (0.012)	0.323*** (0.009)
IV for $\Delta\ln(vm)_{iht-2}$	0.008 (0.006)	−0.003 (0.006)	0.024*** (0.006)	0.021*** (0.006)	0.017*** (0.006)	–
$\Delta\ln(\text{realGDP})_{it}$	3.428*** (0.072)	3.315*** (0.075)	3.014*** (0.083)	3.062*** (0.082)	3.439*** (0.079)	–
$\Delta\ln(\text{realGDP})_{\text{jpn},t}$	−2.434*** (0.185)	−3.238*** (0.198)	−2.640*** (0.221)	−3.239*** (0.217)	−3.501*** (0.217)	–
$\Delta\ln(\text{open})_{it}$	0.391*** (0.045)	0.741*** (0.047)	0.813*** (0.051)	0.790*** (0.051)	0.638*** (0.051)	–
$\Delta\ln(\text{yen/curr}_i)_t$	–	0.265*** (0.025)	0.437*** (0.030)	0.416*** (0.029)	0.268*** (0.027)	–
$\Delta\ln(\text{avg.wage})_{\text{jpn},kt}$	–	–	0.479*** (0.032)	–	–	–
$\Delta\ln(\text{v.add/worker})_{\text{jpn},kt}$	–	–	–	0.362*** (0.025)	–	–
Constant	−0.117*** (0.003)	−0.120*** (0.003)	−0.103*** (0.004)	−0.103*** (0.004)	−0.122*** (0.004)	–
6-digit HS product fixed effects	No	No	No	No	No	Yes
Country *i*-year dummies	No	No	No	No	No	Yes
Observations	141,164	136,583	113,393	113,393	113,393	141,164
Average autocovariance in residuals of order 2 [*z*-stat]	−1.36	−0.73	−0.99	−0.95	−0.99	–
R^2	–	–	–	–	–	0.081

Notes: Subscript i is an importing country, h is a 6-digit HS product, k is a 3-digit ISIC industry, and t is a year. In parentheses are White's heteroskedasticity-consistent standard errors corrected for clusters on the variable defined as the 6-digit HS product and year combination. ***, **, and * denote variables statistically significant at the 1, 5, and 10% levels, respectively.

imposed) facing a Japanese exporter in the sample is 37.13%, which implies a 5.24% average increase in Japanese exports of h to an importing country i. In the next section we investigate whether the magnitude of the trade deflection effect facing Japanese exports varies substantially across some of its particularly important trading partners i as well as important product categories of h.

On the other hand, the imposition of a 1% antidumping duty against the third country i's exporters, but not against Japan, is associated with a 1.269% *reduction* in Japanese exports of that same product h to country i. This is consistent with the idea that when the output produced by firms in country i cannot be sold in the US, but is sold domestically, it depresses (crowds out) country i's imports of the same product from Japan. With the median duty facing a country i's exports of h being 14.84% in this particular sample of data, this translates into an 18.83% average reduction in Japanese exports to a third country, which is also a significant effect.

Finally, and perhaps surprisingly, the US imposition of an MFN safeguard policy also has a strong trade depressing effect. After using the Kennedy (1981) formula to convert the coefficient estimate for the dummy variable to its marginal effects interpretation, the imposition of a US SG on an HS product implies a 67% reduction in Japanese exports of that product to country i.[22]

Next, the coefficient estimates in specification (1) for the macroeconomic and “openness” control variables also have the predicted sign. Since they are not of particular interest to our investigation and are fairly robust across specifications, we will omit a substantive discussion of them here.

In specifications (2) through (4) of Table 3, we sequentially consider additional control variables as one way to check the sensitivity of our results. Overall, the estimates on the policy variables of interest appear robust to changes in model specification. For example, the estimated impacts of the growth in the values of the yen/country i exchange rate, which we add in specification (2), is positive as predicted by the theory. Nevertheless, we do lose a substantial number of observations from the sample when we add in the real exchange rate variable.

More importantly, in specifications (3) and (4) we add industry-level (ISIC 3-digit) control variables (available from 1992–1999). Because our results for the policy variables are robust to the inclusion of industry level controls, we believe that the policy variables are likely capturing the true treatment effect of the policy. Unfortunately, because the industry variables are only available for manufacturing industries between 1992 and 1999, we also lose a number of observations in these specifications. Specification (5) shows that the small changes to the estimates for the policy variables of interest (the slight increase in the size of the AD duty imposition variables and decrease in the statistical significance of the SG policy imposition variable) are most likely due to the loss of observations from years 2000 and 2001. For the case of the safeguard variable in particular, this is likely due to the variation generated by the observations surrounding the 2000 US safeguard policy on circular welded pipe that particularly affected global pipe trade from and into markets such as Korea, Japan and the EU.[23]

[22] While safeguard measures often include exemptions for free-trade partners and developing countries, in our sample, safeguard measures were applied quite broadly to almost all US import sources. The correlation coefficient between changes in US safeguard policy against Japan – the variable in our estimation – and against all other countries was 0.73. In contrast, there was considerably more variation in the application of antidumping duties. The correlation coefficient between changes in US antidumping policy against Japan and against all other countries was only 0.24.

[23] Korea was the largest exporter adversely affected by the US policy, so much so that it contested the measure through a formal WTO trade dispute. It is therefore likely that because of a glut of Korean pipe being retained domestically, Japanese exports of pipe to Korea were depressed in 2000, thus driving the significance of the safeguard results in specifications in which the year 2000 data is included.

At the bottom of Table 3 we also report the z-statistic on the average autocovariance in residuals of order 2 for specifications (1) through (5). For all specifications, we are able to reject the hypothesis of second order autocovariance in the residuals, which leads us to conclude that our Arellano and Bond GMM estimator yields consistent parameter estimates. In all of the specifications reported in (1) through (5) we include two lags of the dependent variable, as we found that inclusion of a second lag improved the fit of the model and yet did not significantly change our parametric estimates.

Finally, specification (6) presents a final robustness check on the results in Table 3. Using the sample from specification (1), we estimate the fixed effects specification of Eq. (6) with the Anderson and Hsiao (1981, 1982) instrumental variables technique, as described in Section 3.4. The results in specification (6) are broadly consistent with those in specifications (1)–(5). The estimated impact of a 0.127% increase in exports in response to a 1% increase in the US antidumping duty against Japan falls within the 95% confidence intervals of the estimated coefficients in specifications (1)–(5), but is not significantly different from zero. We will show in the next section, however, this result appears to be driven by the particular sample of countries and years available in the data set required for the GMM estimation.

To summarize the results of Table 3, we find first that the US imposition of an AD duty against Japan leads to a deflection of Japanese trade to third markets (row 1): Japanese exports to third markets increase by estimates ranging from 0.127% to 0.168% for each 1% increase in the US duty. Second, the US imposition of an AD duty against a third country is associated with the depression of Japanese exports to those third markets (row 2): Japanese exports to third markets fall by 0.870% to 1.292% for each 1% increase in the duty. Third, the US imposition of a broadly applied SG measure against Japan and other exporting countries leads to a depression of Japanese trade to third markets (row 3): Japanese exports to third markets fall by 63% to 70%.

Even though the trade depressing effect of a US safeguard measure is statistically significant, we are concerned about the robustness of this particular result. While there were hundreds of US AD measures imposed over the 1992–2001 period, there were only five US SG investigations which resulted in the imposition of definitive measures (tariffs, quotas or tariff-rate quotas). Even though each of the SG measures may affect more than one 6-digit HS category, we are nevertheless concerned about the relatively few number of safeguard observations in the estimation. This concern is further driven by the fact that some US safeguard measures covered products (e.g., brooms, lamb meat, wheat gluten) which were not of substantial importance to Japanese exporters. Nevertheless, our results with regard to the imposition of a safeguard measure do reject our theoretical model's prediction of two-way trade deflection.

4.2. IV estimates using fixed effects from an expanded sample of data

As described in Section 3.4, an advantage to using the fixed effects model and instrumental variables estimation procedure is that it does not require macroeconomic and industry controls. Thus, we can estimate the model using a significantly larger sample of trade data.[24] Therefore, in Table 4 we provide a set of estimation results using data on Japanese exports to all of the countries listed in Table 1 except the US. When compared to the sample of specification (6) of Table 3, for example, this adds to the estimation sizable import markets such as Taiwan, China, Singapore and the Philippines, in addition to requiring one fewer lag of the dependant variable in the estimation, providing effectively another year (1994) of trade remedy data. Together, these elements add roughly 113,000 observations.

[24] The results presented in Table 4 use country–year dummies and 6-digit HS product fixed effects.

Estimates of the coefficients on the policy variables are consistent with our findings of trade deflection and trade depression reported in Table 3. In Appendix A we formally describe the tests we perform on the strength of our instrumental variables. Because F-tests confirm that all of our instrumental variables are strong instruments, we conclude that our instrumental variables estimates are unbiased.

Specification (7) of Table 4 presents the baseline IV specification for comparison with the results of Table 3.[25] The estimates for trade deflection and trade depression are statistically significant on this larger sample of data. The primary change in results from Table 3 relates to the size of the coefficient estimate on the trade *depression* effect of an antidumping duty against country i, as it falls from -1.271 in specification (1) to -0.281 in specification (7).

Nevertheless, to better interpret and compare the magnitude of the estimates, consider Table 5. The first column presents the median duty, conditional on a duty being applied, for the sample of 141,164 observations used in specifications (1)–(6). The second column presents the median duty, conditional on a duty being applied, for the sample of 254,074 observations used in specifications (7)–(10). The third column quantifies the effect of imposing a typical duty, in this case, the conditional median[26] using the coefficient estimated from specification (1). The fourth and fifth columns similarly quantify the effect of imposing a typical duty using the coefficient estimates from specifications (7) and (8). First, comparing our estimates from specifications (1) and (7), we see that the magnitudes of the trade deflection effect in the two samples, using two different econometric specifications, is similar. In our GMM specification (1), imposing the conditional median duty is associated with a roughly 5% increase in Japanese exports of that product to third countries while for our IV specification (7), imposing the conditional median duty is associated with a roughly 7% increase in Japanese exports. There is, however, a noticeable difference in the magnitude of the trade depression effect between specifications (1) and (7). The depression effect associated with the conditional median in the GMM specification (1) of a 19% fall in Japanese exports to country i is considerably larger than that in the IV specification (7) of a 5% fall in Japanese exports. We believe this difference in the magnitude of trade depression is likely due to differences in the underlying sample of data. In particular, the GMM sample requiring macroeconomic data does not contain observations for a number of Japan's sizable export markets, including China, Taiwan, Singapore and the Philippines. Furthermore, the GMM procedure also requires an additional lag of the trade data, and the result for those specifications is to lose all useful trade remedy variation taking place in 1994.

Thus, in specification (8) of Table 4 we explore the question of whether there are substantial differences in trade deflection and trade depression effects for Japanese exports across importing country markets. This specification is estimated on the identical sample of data as specification (7), but in this column we present estimates where we interact the antidumping variables of interest with a number of importing country indicators, examining the variation across some of Japan's important export markets. This approach yields strong evidence of trade deflection, for example, associated with Japan's exports to both the EU and Korea, which is quite intuitive, given that they are the two largest destination markets for Japanese exports in the sample (see again Table 1). These two countries are thus likely to be the "next best" alternative markets for Japanese exports that get shut out of the US because of a US trade policy. Table 5 further illustrates the significance of the economic magnitude of Japan's trade deflection to the EU and Korea, as the median duty imposed against products which the Japanese export to these countries results in a 16% increase in Japanese exports to

[25] Note that specification (10) presents the biased OLS estimates as a benchmark for comparison.

[26] In all our samples, the mean duty conditional on a duty being imposed is larger than the median. Thus, when we use the conditional mean duty as a typical duty to quantify the magnitudes of trade deflection and trade depression, the results are slightly (1–5 percentage points) larger.

Table 4
IV estimation using fixed effects: results for Japanese exports, 1994–2001

Explanatory variables	Dependent variable: $\Delta\ln(vm_{iht})$			
	Full sample IV specification (7)	IV by country (8)	IV by commodity (9)	Full sample OLS specification (10)
AD duty changes against Japan				
ΔAD duty$_{\mathrm{jpn},ht}$	0.190*** (0.061)	–	–	0.105* (0.059)
ΔAD duty$_{\mathrm{jpn},ht}$ for exports sent to EU	–	0.390* (0.222)	–	–
ΔAD duty$_{\mathrm{jpn},ht}$ for exports sent to Korea	–	0.578** (0.242)	–	–
ΔAD duty$_{\mathrm{jpn},ht}$ for exports sent to China	–	−0.326 (0.203)	–	–
ΔAD duty$_{\mathrm{jpn},ht}$ for exports sent to India	–	0.238 (0.454)	–	–
ΔAD duty$_{\mathrm{pn},ht}$ for exports sent to other	–	0.198*** (0.070)	–	–
ΔAD duty$_{\mathrm{jpn},ht}$ for exports of steel	–	–	0.157 (0.100)	–
ΔAD duty$_{\mathrm{jpn},ht}$ for exports of non-steel	–	–	0.204*** (0.079)	–
AD duty changes against country i				
ΔAD duty$_{iht}$	−0.281*** (0.101)	–	–	−0.242** (0.096)
ΔAD duty$_{i,ht}$ for exports sent to EU	–	−0.812 (0.915)	–	–
ΔAD duty$_{i,ht}$ for exports sent to Korea	–	−1.000 (1.830)	–	–
ΔAD duty$_{i,ht}$ for exports sent to China	–	−0.249** (0.120)	–	–
ΔAD duty$_{i,ht}$ for exports sent to India	–	−1.118***(0.308)	–	–
ΔAD duty$_{\mathrm{pn},ht}$ for exports sent to other	–	0.145 (0.243)	–	–
ΔAD duty$_{i,ht}$ for exports of steel	–	–	−0.077 (0.160)	–
ΔAD duty$_{i,ht}$ for exports of non-steel	–	–	−0.413*** (0.130)	–
SG policy changes				
ΔSG policy$_{ht}$	−0.807*** (0.208)	−0.805*** (0.208)	–	−0.547*** (0.198)
ΔSG policy$_{ht}$ for exports of steel	–	–	−0.868*** (0.240)	–
ΔSG policy$_{ht}$ for exports of non-steel	–	–	−0.621 (0.415)	–
Other control variables				
IV for $\Delta\ln(vm)_{iht-1}$	0.306*** (0.006)	0.306*** (0.006)	0.306*** (0.006)	–
$\Delta\ln(vm)_{iht-1}$	–	–	–	−0.308*** (0.002)
Constant	0.339*** (0.035)	0.338*** (0.035)	0.339*** (0.035)	−0.682 (0.923)
6-digit HS product fixed effects	Yes	Yes	Yes	Yes
Country *i*-year dummies	Yes	Yes	Yes	Yes
Observations	254,074	254,074	254,074	254,074
R^2	0.065	0.065	0.065	0.147
F-statistic	27370.6	27497.1	28025.0	–

Table 5
Magnitude of trade deflection and trade depression estimates

			Percent increase in affected Japanese product-level exports associated with subsample's median duty, conditional on a duty being imposed		
	Subsample's median conditional duty [141,164 observations]	Subsample's median conditional duty [254,074 observations]	GMM specification (1)	IV specification (7)	IV specification (8)
Trade deflection					
ΔAD duty$_{jpn,ht}$	37.13	–	5.24*	–	–
ΔAD duty$_{jpn,ht}$	–	37.13	–	7.05***	–
ΔAD duty$_{jpn,ht}$ for exports sent to EU	–	41.72	–	–	16.27*
ΔAD duty$_{jpn,ht}$ for exports sent to Korea	–	34.83	–	–	20.13**
ΔAD duty$_{jpn,ht}$ for exports sent to China	–	37.13	–	–	−12.10
ΔAD duty$_{jpn,ht}$ for exports sent to India	–	32.52	–	–	7.74
ΔAD duty$_{jpn,ht}$ for exports sent to other	–	37.13	–	–	7.35***
Trade depression					
ΔAD duty$_{iht}$	14.84	–	−18.83***	–	–
ΔAD duty$_{iht}$	–	16.26	–	−4.57***	–
ΔAD duty$_{iht}$ for exports sent to EU	–	10.43	–	–	−8.45
ΔAD duty$_{iht}$ for exports sent to Korea	–	3.08	–	–	−3.08
ΔAD duty$_{iht}$ for exports sent to China	–	121.46	–	–	−30.24**
ΔAD duty$_{iht}$ for exports sent to India	–	72.49	–	–	−81.04***
ΔAD duty$_{iht}$ for exports sent to other	–	20.28	–	–	2.94

Notes: Subscript i is an importing country, h is a 6-digit HS product, and t is a year. ***, **, and * denotes underlying coefficient estimate was statistically significant at the 1, 5, and 10% levels, respectively. Specification (1) uses the coefficient estimates from Table 3 and specifications (7) and (8) use the coefficient estimates from Table 4.

the EU and a 20% increase in Japanese exports to Korea. This is considerably larger than the roughly 7% increase in Japanese exports to countries other than the EU, Korea, China, and India.

Turning to the country-specificity of trade depression, specification (8) of Table 4 indicates that US antidumping duties on third countries are associated with statistically significant reduction in Japanese exports to China and India. Table 5 illustrates the economic significance of these results as well, quantifying an 81% decrease in Japanese exports to India and a 30%

Notes to Table 4
Subscript i is an importing country, h is a 6-digit HS product, and t is a year. In parentheses are White's heteroskedasticity-consistent standard errors corrected for clusters on the variable as the 6-digit HS product and year combination. ***, **, and * denote variables statistically significant at the 1, 5, and 10% levels, respectively. F-statistic is for the weak instrument test described in Appendix A.

decrease in Japanese exports to China associated with the imposition of median US antidumping duties against each of these two countries. We believe that the magnitude and significance of trade depression for India and China in particular may be related to three phenomena. First, the second column of Table 5 indicates that both of these countries face extremely high US antidumping duties: the median duty imposed against Chinese exports to the US was 121% while the median duty against Indian exports was 72%. It seems likely that duties facing these two countries are frequently prohibitive, which would create a severe glut of the affected products in the Chinese and Indian markets which could crowd out imports from Japan and lead to a sizable amount of trade depression. Second, unlike a number of other importing countries in the sample, China and India are also frequent targets of US antidumping activity. And finally, even beyond a higher frequency of being targeted, both of these countries are also less frequently targeted alongside Japan in a multi-country US AD investigation over the same product (i.e., relative to the EU and Korea, whose exporters have a higher frequency of being alongside Japan in a multi-country US AD investigations over the same product). This variation likely allows for more precise estimation of the trade depressing effect associated with a US antidumping duty being applied on third countries such as China and India alone (i.e., not simultaneously with Japan). This is also a potential explanation for the imprecisely estimated trade depressing effects of US antidumping toward the EU and Korea in specification (8) of Table 4.

4.3. IV estimates for steel versus non-steel products

Another question to consider is whether the AD or SG measures associated with the US steel industry are particularly important in our results, given that this industry is the most frequent user of US trade remedies.[27] To address this issue, in specification (9) of Table 4 we separate out the estimated policy effects for steel and non-steel products by interacting each policy variable of interest with an indicator for whether the underlying 6-digit HS product was a steel (HS chapter 72 or 73) or non-steel product. With the exception of the estimates for the SG policy (which again are tested on a relatively small number of policy actions), the estimates suggest that the trade deflection and trade depression results may be even stronger for non-steel products than the steel products that have traditionally been the most active targets of US trade remedy laws. This is particularly important, given the likelihood that any future growth in use of US trade remedies is likely to come from non-steel industries as they learn from the steel industry's experience. Thus this table might suggest that future use of trade remedies may lead to even more trade deflection and trade depression.

4.4. Specification tests

As a final check on our results, we conduct a specification test on our GMM (5) and IV-fixed effects (6) econometric models. These results are reported in Table 6. The thought experiment is similar to that conducted in the labor literature, beginning with Ashenfelter (1978), on the evaluation of training programs for unemployed workers. As applied to our context, we investigate whether the imposed policy (i.e., an AD duty) can be used to predict the change in the dependent variable that occurred before the policy was imposed. Specifically, we might be concerned that a Japanese, product-level cost shock in $t-1$ led to an increase in Japanese exports to US and non-US markets in $t-1$, and thus that

[27] For example, for the 1992–2001 period, over 50% of US AD investigations that resulted in duties affected steel imports.

Table 6
Specification test: Do US trade remedies predict Japanese export growth to third countries before they're imposed?

Explanatory variables	Dependent variable: $\Delta\ln(vm_{iht-1})$ [lagged Japanese export growth of product h to importer i]		
	Test of specification (1)	Test of specification (2)	Test of specification (7)
AD duty changes against Japan			
ΔAD duty$_{\text{jpn},ht}$	−0.124 (0.109)	−0.152 (0.108)	−0.053 (0.072)
AD duty changes against country i			
ΔAD duty$_{iht}$	0.391 (0.349)	0.396 (0.345)	0.148 (0.144)
SG policy changes			
ΔSG policy$_{ht}$	−0.016 (0.118)	−0.006 (0.117)	−0.099* (0.055)
Other control variables			
IV for $\Delta\ln(vm)_{iht-2}$	0.277*** (0.012)	0.245*** (0.012)	0.306*** (0.006)
IV for $\Delta\ln(vm)_{iht-3}$	0.009 (0.006)	−0.002 (0.006)	–
$\Delta\ln(\text{realGDP})_{it-1}$	3.425*** (0.072)	3.310*** (0.075)	–
$\Delta\ln(\text{realGDP})_{\text{jpn},t-1}$	−2.412*** (0.185)	−3.222***(0.198)	–
$\Delta\ln(\text{open})_{it-1}$	0.395** (0.045)	0.747*** (0.047)	–
$\Delta\ln(\text{yen/curr}_i)_{t-1}$	–	0.267*** (0.045)	–
Constant	−0.117*** (0.003)	−0.120*** (0.003)	0.339*** (0.035)
6-digit HS product fixed effects	No	No	Yes
Country *i*-year dummies	No	No	Yes
Observations	141,162	136,581	254,070
Average autocovariance in residuals of order 2 [*z*-statistics]	−1.35	−0.72	–
R^2	–	–	0.065

Notes: subscript *i* is an importing country, *h* is a 6-digit HS product, and *t* is a year. In parentheses are White's heteroskedasticity-consistent standard errors corrected for clusters on the variable defined as the 6-digit HS product and year combination. ***, **, and * denote variables statistically significant at the 1, 5, and 10% levels, respectively.

the US policy response in t could actually be used to predict the $t-1$ export growth to importing countries *i*. If this were the case, our model could be misspecified and what we claim to be trade deflection might just be an increase in Japanese exports associated with a favorable cost shock to a particular product exported to many different markets.

To investigate this question we therefore regress lagged ($t-1$) product-level Japanese export growth to country *i* on period *t* US policy changes and other explanatory variables.[28] If the coefficient on a US policy change in *t* were positive and statistically significant, the imposition of the US AD policy could be interpreted as a predictor of higher than normal Japanese export growth of that product to all markets in $t-1$ which, in turn, might have been due to a cost shock. Similarly, if US imposition of an AD duty against country *i* in *t* were a statistically significant predictor of a reduction in Japanese exports to *i* in $t-1$, this could suggest that a product-level cost shock in the importing country was behind what we have described as "trade depression." Nevertheless, in all specifications in Table 6, the coefficients on changes in US AD policies against Japan and country *i* are not statistically significant. Our results indicate that the model is correctly specified and that the policy variables are measuring the true treatment effect of a policy change.

[28] Due to the small number of changes in safeguards policy, there were insufficient observations to perform this specification test using safeguards.

5. Conclusion

This paper empirically examines whether a country's use of an import-restricting trade policy distorts a foreign country's exports to third markets. To investigate this question we match data on US use of antidumping and safeguard trade remedies over the 1992–2001 period to Japanese product-level exports to third countries. We find evidence that US trade remedies both deflect and depress Japanese exports. The median antidumping duty against Japan leads to a 5–7% average increase in Japanese exports to a non-US trading partner. When the median US antidumping duty is imposed against a third country's exporters, Japanese exports in the same product to that third country decrease by an average of 5–19%. Finally, when faced with a US safeguard measure, Japanese exports to third countries fall by somewhere between 55% and 70%. Our results on the "deflection" and "depression" of Japanese exports vary substantially (and in intuitively appealing ways) across importing countries, and the estimated impact appears stronger for non-steel relative to steel products.

There are some limitations of our results and approach. First, we have focused on the export response of only one US trading partner. An open research question is whether US trade policy similarly distorts the exports of other trading partners, including developing countries. We speculate, for example, that the ability of developing countries to deflect trade may be more limited than that of a country like Japan. Furthermore, we are less confident in our results regarding the impact of safeguard policies, as there are relatively few safeguard observations in our dataset.

Nevertheless, our results have implications for the empirical literature on the impact of trade policy decisions made by "large" countries, defined as those that are able to affect exporters' prices. For example, Chang and Winters (2002) use similarly disaggregated, product level data on unit values and tariffs for Brazil and its trading partners and find that the creation of MERCOSUR was accompanied with a substantial decline in the prices of non-member exporters to the region. While we do not test whether any of the countries in our analysis are "large" in the sense of their ability to affect the prices of foreign exporters, we provide evidence that the US's trade policy decisions do impact the export behavior of a particularly important trading partner. Finally, we speculate that the results of this paper suggest an additional explanation for the proliferation of antidumping laws around the world (Miranda et al., 1998; Prusa, 2001) that has not previously been investigated. Much of the prior literature commenting on this proliferation has focused on the retaliation argument: countries adopt trade remedy laws in order to establish a credible retaliatory threat that will discourage foreign trade remedies targeted against their exporters (Prusa and Skeath, 2002; Blonigen and Bown, 2003). Our results indicate that the imposition of a US trade remedy can lead to a substantial export surge to a third country's market. This third country may therefore face pressure of its own to respond with a trade remedy. Therefore, US actions may induce trade policy actions by third country importers in addition to (and that is separate from) retaliation-based trade policy actions. While we do not test here for the formal link between US trade policy actions and responses by the governments of third countries facing deflected trade, our results that associate substantial export surges with US trade policy changes suggests an additional explanatory factor that should be an area of future research.

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Table A-1
Testing instrument quality

Explanatory variables	Dependent variable: $\Delta\ln(vm_{iht})$			
	First-stage unrestricted regression based on specification (7)	First-stage restricted regression based on specification (7)	First-stage unrestricted regression based on specification (9)	First-stage restricted regression based on specification (9)
AD duty against Japan				
ΔAD Duty$_{\mathrm{jpn},ht}$	−0.131** (0.060)	−0.137** (0.063)	–	–
ΔAD duty$_{\mathrm{jpn},ht}$ for exports of steel	–	–	−0.053 (0.096)	−0.043 (0.101)
ΔAD duty$_{\mathrm{jpn},ht}$ for exports of non-steel	–	–	−0.175** (0.076)	−0.189** (0.080)
AD duty changes against country i				
ΔAD duty$_{iht}$	0.034 (0.097)	0.069 (0.102)	–	–
ΔAD duty$_{\mathrm{i},ht}$ for exports of steel	–	–	−0.151 (0.154)	−0.198 (0.162)
ΔAD duty$_{\mathrm{i},ht}$ for exports of non-steel	–	–	0.150 (0.125)	0.237* (0.131)
SG policy changes				
ΔSG policy$_{ht}$	0.397** (0.200)	0.423** (0.211)	–	–
ΔSG policy$_{ht}$ for exports of steel	–	–	0.225 (0.231)	0.211 (0.243)
ΔSG policy$_{ht}$ for exports of non-steel	–	–	0.909** (0.400)	1.065** (0.421)
Other control variables				
$\Delta\ln(vm)_{iht-1}$	−0.210*** (0.001)	–	−0.210*** (0.001)	–
6-digit HS product fixed effects	Yes	Yes	Yes	Yes
Country *i*-year dummies	Yes	Yes	Yes	Yes
Observations	254,598	254,598	254,598	254,598
R^2	0.152	0.059	0.154	0.059

Notes: Subscript *i* is an importing country, *h* is a 6-digit HS product, and *t* is a year. In parentheses are White's heteroskedasticity-consistent standard errors corrected for clusters on the variable defined as the 6-digit HS product and year combination. ***, **, and * denote variables statistically significant at the 1, 5, and 10% levels, respectively.

Appendix A. Instrument tests

There are two potential problems with the IV estimator used in estimating Eq. (6); bias associated with the use of a weak instrument and bias associated with correlation in measurement error.

First, in a dynamic panel model, if the autoregressive coefficient on imports is sufficiently large, then the lagged level of imports, $\ln(vm_{iht-2})$, will be a weak instrument for the lagged difference, $\Delta\ln(vm_{iht-1})$ (Blundell and Bond, 1998). In this case, the bias of the IV estimator in a small sample is large (Nelson and Startz, 1990). To test the quality of two instruments, $\ln(vm_{iht-2})$ and $\ln(vm_{iht-3})$, the following first-stage model was estimated using each instrument for each of the specifications presented in Table 4,

$$\Delta\ln(vm_{iht-1}) = \mu_h + \chi_{it} + \eta_1^{'}\Delta\tau_{ht} + \eta_2^{'}\Delta\tau_{iht} + \eta_3^{'}\ln(vm_{iht-2}) + \Delta\epsilon_{iht-1}, \tag{7}$$

where $\ln(vm_{iht-3})$ was substituted for $\ln(vm_{iht-2})$ in some specifications. As a restricted regression, Eq. (7) was estimated under the assumption that η_3 is equal to zero. Table A-1 reports results using the parameters of specification (7) and specification (9) in Table 4. Results for all other specifications are similar. For the model based on the parameters in specification (7), the F-statistic of 27,370 is far larger than the 99% critical χ^2 (1) of 6.63. Likewise, for the model based on the parameters in specification (9), the F-statistic of 28,025 is larger than the 99% critical χ^2 (1) of 6.63. In all specifications, we find that $\ln(vm_{iht-2})$ and $\ln(vm_{iht-3})$ are strong instruments for $\Delta\ln(vm_{iht-1})$ and conclude the IV approach is appropriate for our problem.

Second, consider the use of the second lag of the logged level, $\ln(vm_{iht-2})$, as an instrument for $\Delta\ln(vm_{iht-1})$. If there is measurement error in $\ln(vm_{iht})$, then measurement error in the regressor, $\Delta\ln(vm_{iht-1})$, will be correlated with measurement error in the instrument, $\ln(vm_{iht-2})$, and the IV estimator will be biased. An alternative IV, the third lag of the logged level, $\ln(vm_{iht-3})$, has the advantage that its measurement error will not be correlated with measurement error in the regressor. The disadvantage of this instrument is that it further shortens an already short panel. Our approach is to estimate Eq. (6) using each of these instruments for every IV specification reported in Table 4. By necessity, this requires using the small sample that obtains when we use the third lag of the level as the instrument. We find that the coefficient estimates are robust to the choice of instrument, suggesting that measurement error in $\ln(vm_{iht})$ is not a significant problem and the use of $\ln(vm_{iht-2})$ as an instrument for $\Delta\ln(vm_{iht-1})$ is appropriate.

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