US Policy Spillover(?) – China's Accession to the WTO and Rising Exports to the EU

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This version: December 11, 2015

Abstract

The paper explores a novel mechanism to explain the rise China's manufacturing exports after WTO accession. Building on a literature which describes export market entry by firms in low-wage countries, a theoretical model is developed to analyze the implications of reduced trade policy uncertainty in a major market. A global (non-destination specific) and separable component of exporting fixed costs facilitates transmission of bilateral policy changes to the firms' multilateral export performace. The empirical analysis exploits the removal of US tariff uncertainty in conjunction with China's WTO accession, and examines its effect on China's exports to the EU. The results reveal that: (i) the structure of China's export boom to the EU conforms to the pattern of US tariff uncertainty; (ii) the adjustment takes place at the extensive margin, (i.e., a good is exported to more destinations); and (iii) the effect phases out after a few years. These findings bear important implications for the scope of international policy negotiations and provide suggestive evidence on the nature of fixed costs manufacturing firms in low-wage countries must overcome.

JEL-Classification: F13, F14, D84, O24
Keywords: Exports, China, WTO, Policy Uncertainty, Fixed Costs, Spillover

1 Introduction

When China entered the World Trade Organization (WTO), in December 2001, industrialized countries saw increasing amounts of Chinese goods flowing into their domestic markets. Numerous studies have been analyzing the consequences of the increased competition on domestic firms and workers.[1] Investigating such questions requires understanding the causes of China's export boom. Reduced import duties on Chinese products, the most intuitive explanation, does not appear to be the main reason, because tariffs often corresponded to the most-favored-nations (MFN) rates even before WTO entry. China's liberalization of foreign investment regulations has certainly contributed to rising exports, but their structural pattern across manufacturing industries is difficult to measure or infer. Other changes in Chinese policies, such as reduced tariffs on imported intermediate inputs (Feng et al., 2012), or the gradual removal of export controls (Bai et al., 2015), could be identified as meaningful triggers of (transitional) export growth in specific industries. Reduced trade policy uncertainty (TPU) has recently become another promising approach to assess the effect of China's WTO entry. In fact, it could be shown that the removal of uncertainty regarding US tariff policies towards China explains well timing and structure of China's export boom to the US (Handley and Limão, 2013; Feng et al., 2014), as well as sectoral patterns of declining US manufacturing employment (Pierce and Schott, 2013).

Yet, the assessments of reduced tariff uncertainty have been limited to the bilateral trade relations among the involved countries. While this is naturally the first step of evaluating trade-creating effects of TPU reductions, stylized facts on exporters in low-wage countries, and particularly China, motivate investigations beyond this. The sizeable contribution of processing trade to China's overall export growth (Amiti and Freund, 2010) and the differential exporting behavior of private domestic and foreign-invested enterprises in China (Manova and Zhang, 2009) justify speculations that export market entry of these firms follows different

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[1]See, for instance, Bernard et al. (2006); Bloom et al. (2011); Autor et al. (2013a,b); Dauth et al. (2014); Utar (2014, 2015) for the most recent contributions.

rules than those derived from firm-level studies in high-income regions.[2] For instance, firms in emerging countries are found to upgrade their products or hire new staff before becoming exporters (Iacovone and Javorcik, 2012; Molina and Muendler, 2013). This suggests presence of sunk investments that are not necessarily specific to a particular export destination.

The present paper analyzes the implications reduced TPU in a major market may have on exports to third countries. Feng et al. (2014) and Handley and Limão (2013) analyzed China's US exports in the bilateral setting, suggesting that monopolistic heterogenous firms are selected into exporting, due to lower expected tariffs. However, their theoretical analyses do not yield any direct or indirect effects on exporters' performance in third countries. Different from their models, this paper considers a fixed costs structure which resembles characteristics of a services trade model (Hanson and Xiang, 2011). When exporters face fixed costs, which are separable into a destination-specific and a product-specific (global) component, a reduction of tariff uncertainty in the US may have stimulated China's multilateral exports, as well. In fact, the theoretical analysis reveals that a removal of bilateral tariff uncertainty exhibits a spillover on exports to third countries, which results from a redistribution of the global fixed cost burden across sources of income. The effect is expectedly larger when the policy-making country is large, while the selection of new firms into exporting implies an adjustment the extensive margin.

To test the predictions, the paper uses disaggregated data on Chinese manufacturing exports to the European Union (EU-15) in the period 1995-2005. The impact of the US policy change towards China is evaluated using a difference-in-difference (DID) strategy, exploiting the cross-product differences of the US "tariff threat" under TPU, at the HS6 product level, and comparing years before and after China's WTO entry. The EU-15 are chosen as destinations, because they represent a large integrated high-income market with common customs regulations. Importantly, EU tariff policies towards China were less descretionary than in the US and did not change upon WTO entry.

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[2]A comprehensive characterization of French firms' exporting behaviour is presented by Eaton et al. (2004, 2011).

The results reveal a positive impact of the US policy change on Chinese exports to the EU, which is robust to the inclusion of variables that control for potentially confounding effects and spurious correlation. The findings do not appear to be driven by exports to an individual European destination, or by the EU as a whole, as the pattern can be confirmed also for Chinese exports to Japan. The estimated coefficients are in reasonable orders of magnitude, considering model parameter values used in the literature, and it can be confirmed that trade increases through initiation of new trade relationships. Further analyses suggest that the effect phases out after a few years, which suggests that new exporters do not systematically outperform former exporters. Suggestive evidence is revealed in support of a redistribution of global fixed costs, whereas it seems that the goods increasingly exported are different from those traded within the East Asian production networks. Affected products are less likely to be differentiated and more likely to be utilized as low-intensity inputs in a wide range of industries.

An important insight of this paper is that bilateral trade policies are not without consequences for third countries, especially when large economies are involved. This extends the scope of the trade creating effects that have been recently identified in bilateral studies (Handley, 2012; Handley and Limão, 2012, 2013). However, the type of countries making the policy and those receiving the policy seem to matter. The proposed transmission channel, a global fixed cost component, generalizes the findings of Hanson and Xiang (2011), who developed a model for services exports. Although global fixed costs seem to be at odds with evidence from manufacturing firm studies in high-income regions (e.g. Eaton et al., 2011), a potential reason for their relevance, in the presented case, is the way how Chinese firms participate in international trade (Amiti and Freund, 2010; Manova and Zhang, 2009). They carry out production, and thus export labor services, while their clients focus on developing and placing goods on the market. Hence, the product-specific fixed cost is paid to meet certain criteria that attract orders. Country-specific fixed costs remain an important feature of manufacturing trade, however, because they influence coordination and transportation

costs of the goods. The suggested segmentation of China's exports into trade within specialized production networks and large-scale supply of standardized goods to distant locations conforms to theories where the latter mode of exporting is more exposed to international wage competition (Thesmar and Thoenig, 2002; Holmes and Stevens, 2014). Altogether, the export patterns for a less dynamic group of manufacturing production are presented. Nevertheless, this group is important, particularly in developing countries, where labor cost advantage seems to be one prerequisite for participation in international trade.

The structure of the paper is as follows. Section 2 introduces the theoretical model, which highlights the differential implications of reduced tariff uncertainty obtained from alternative fixed cost structures. Numerical examples provide intuition on the transmission of bilateral policies to multilateral exports, so that testable predictions can be formulated. Section 3 describes the trade policy environments Chinese exporters faced in the EU and in the US, which will be the setting to evaluate the hypotheses derived from the theory. It decribes further the empirical strategy and the data used for estimation. Section 4 presents and discusses the main results, robustness checks, and further findings. Section 5 concludes.

2   Modelling the Policy Spillover

Recent work on the effect of TPU reductions focused on bilateral trade relationships, modelling fixed costs as a single term that varies in a product-destination dimension. Productivity is heterogeneous and Pareto-distributed across monopolisitc firms (Melitz, 2003). The first part of this section illustrates the mechansim that yields higher bilateral exports after a reduction of bilateral tariff uncertainty, within such frameworks (Handley, 2012; Handley and Limão, 2013; Feng et al., 2014). The second part alters the fixed costs structure and analyzes the effect bilateral tariff uncertainty has on multilateral exports.

2.1 Baseline Model

2.1.1 Setup

Demand. Monopolistic firms j consider demand of utility-maximizing consumers. Consumers in country n allocate a fraction $0 < \mu < 1$ of their expenditures on product J across foreign varieties X_J. The rest is spent on a domestic numéraire, 0_J. X_J is defined as a CES aggregator over available varieties $j \in \Omega_{Jn}$ so that demand is given by

$$X_{Jn} = \left(\int_{j \in \Omega_{Jn}} x_j^\epsilon dj \right)^{1/\epsilon}. \tag{1}$$

The elasticity of substitution is stated in the exponents, $\sigma \equiv 1/(1-\epsilon) > 1$. Total expenditure on differentiated goods, E_{Jn}, the price for a variety, p_{jn}, and the aggregate price index, $P_{Jn} \equiv \left[\int_{j \in \Omega_{Jn}} p_{jn}^{1-\sigma} dj \right]^{1/(1-\sigma)}$, determine the demand for variety j in country n:

$$x_{jn} = \frac{E_{Jn}}{P_{Jn}} \left(\frac{p_{jn}}{P_{Jn}} \right)^{-\sigma} \tag{2}$$

Supply. Each monopolistic firm produces only one variety and charges a mark-up over its marginal costs. The price consumers in destination n have to pay for variety j is determined by the firm's productivity parameter, $\varphi(j)$, wages in the exporting country, w, the costs of shipping the good to country n, $d_{Jn} \geq 1$, and by the tariff rate, $\tau_{Jn} \geq 1$,

$$p_{jn} = \left(\frac{\sigma}{\sigma - 1} \right) \frac{w}{\varphi_j} d_{Jn} \tau_{Jn}. \tag{3}$$

The only variety-specific component is the productivity of firm j.

The operating profit function is given by $\pi_{jn}^{op} = (\tilde{p}_j - c_j) x_{jn}$, where $\tilde{p}$ corresponds to the market price discounted by destination-specific costs, i.e. $\tilde{p}_j \equiv p_{jn}/(d_{Jn}\tau_{Jn})$, and $c_j = w/\varphi_j$ denotes marginal production costs. Substituting (2) and (3) into the operating profits function gives

$$\pi_{jn}^{op} = \tau_{Jn}^{-\sigma} d_{Jn}^{-\sigma} \varphi_j^{\sigma-1} A_{Jn}.$$

The term $A_{Jn} \equiv (1 - \epsilon)E_{Jn}(w/P_{Jn}\epsilon)^{1-\sigma}$ summarizes aggregate conditions. Importantly, tariffs and transportation costs are fully external. They are collected by the destination country's customs office and by firms carrying out logistics and supply. This implies that they are not part of the firm's value added or net revenues.[3]

Exporting Fixed Costs and Firm Entry. Exporting requires additional sunk costs. They are independent of the quantity a firm sells, but may be high enough to prevent a firm from selling its good abroad. Assuming that these costs are inseparably specific to product-destination pairs, yields the net exporting profits:

$$\pi_{jn} = \tau_{Jn}^{-\sigma} d_{Jn}^{-\sigma} \varphi_j^{\sigma-1} A_{Jn} - f_{Jn}. \tag{4}$$

For any firm j, profitable exports require $\tau_{jn}^{op} \geq f_{Jn}$. The marginal firm, which is indifferent between exporting and not exporting has $\tau_{jn}^{op} = f_{Jn}$, and productivity

$$\varphi_{Jn}^* = \tau_{Jn}^{\frac{\sigma}{\sigma-1}} d_{Jn}^{\frac{\sigma}{\sigma-1}} \left[\frac{f_{Jn}}{A_{Jn}} \right]^{\frac{1}{\sigma-1}}. \tag{5}$$

Equation (5) states the zero-profit cutoff (ZPC) productivity level of the least productive exporter. He must have higher productivity, if applied tariffs, τ, shipping costs, d, or fixed costs, f, are high, or if demand, E, or prices, P, in the destination are low.

2.1.2 Tariff Uncertainty

The analysis of tariff uncertainty considers that τ_{Jn}^s is different under different policy regimes, exporters may face in the destination: $s = \{p, np\}$. If the importing country grants preferen-

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[3]This assumption is very similar to Handley and Limão (2013). It is used here to emphasize that firms j are pure producers. They are not involved into destination-specific product development or marketing activities. Qualitatively, this assumption has no impact on the results.

tial market access ($s = p$), the tariff is lower than with non-preferential access ($s = np$), i.e., $\tau^p \leq \tau^{np}$. In the context of this paper, interest lies on the removal of uncertainty regarding the application of preferential tariffs. In presence of uncertainty, firms do not know whether preferential rates will continue to apply in the future. They form expectations on the probability, $0 \leq \delta \leq 1$, that non-preferential tariffs will be levied. Without further characterization of the formation of expectations, expected tariff can be written as the weighted geometric mean of the two scenarios, $\tau^E = (\tau^{np})^\delta (\tau^p)^{(1-\delta)}$.[4] It implies that $\tau^p \leq \tau^E \leq \tau^{np}$. Equation (5) can be rewritten as

$$\varphi_{Jn}^* = (\tau_{Jn}^E)^{\frac{\sigma}{\sigma-1}} d_{Jn}^{\frac{\sigma}{\sigma-1}} \left[\frac{f_{Jn}}{A_{Jn}} \right]^{\frac{1}{\sigma-1}} \tag{6}$$

and illustrates the inhibiting effect tariff uncertainty has on entry of firms into a foreign market.

Lemma 1 *If $\tau_J^p < \tau_J^{np}$, a removal of tariff uncertainty implies a reduction of the ZPC productivity level, φ^*, firms must achieve to export profitably.*

2.1.3 Product-level Predictions for Bilateral Trade

Lemma 1 states the theoretical motivation for an assessment of bilateral TPU reductions (e.g. Handley and Limão, 2013; Feng et al., 2014). Due to inseparable fixed costs, f_{Jn}, it is limited to the relationship between the exporting country and the policy-making destination n. Before analyzing an alternative fixed-costs structure, the product-level predictions of this model are persented.

Firms' export revenues, $r_{jn} = \tilde{p}_{jn} x_{jn} = \sigma \pi_{jn}^{op}$, can be aggregated over product J:

$$R_{Jn} = a_{Jn}\sigma \left(\int_{j \in \Omega_{Jn}} \varphi_j^{\sigma-1} dj \right). \tag{7}$$

The term $a_{Jn} \equiv (\tau_{Jn}^E)^{-\sigma} d_{Jn}^{-\sigma} A_{Jn}$ summarizes all product- and destination-specific variables,

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[4]See Handley and Limão (2013) and Feng et al. (2014) for more explicit characterizations of expected tariffs. Their empirical applications, however, both compare scenarios analogous to moving from $\delta > 0$ to $\delta = 0$, so that the notation used here should be sufficient to illustrate the policy spillover.

so that it can be multiplied with the aggregated productivity of the exporting firms. The expression in parentheses is equivalent to multiplying the total number of firms that produce varieties of J, M_J, with the fraction of firms residing at or above the ZPC,

$$R_{Jn} = a_{Jn}\sigma M_J \left(\int_{\varphi^*_{Jn}} \varphi^{\sigma-1} dG_J(\varphi) \right).\tag{8}$$

If productivity levels across firms are Pareto distributed, the probability that a random draw from this distribution exceeds its lower bound, φ_L, equals $G(\varphi) = \left(\frac{\varphi_L}{\varphi}\right)^k$.[5] Integrating (8) with $G(\varphi)$ gives

$$R_{Jn} = a_{Jn} \left(\frac{1}{\varphi^*} \right)^{k-\sigma+1} \alpha_J,\tag{9}$$

where $\alpha_J \equiv \sigma M_J \varphi_L^k \frac{k}{k-\sigma+1}$ represents a product-specific intercept, and a gravity equation is obtained by plugging (6) into (9) and taking logs:

$$\begin{aligned}
\ln R_{Jn} = \quad &- \frac{\sigma k}{\sigma-1} \ln \tau^p_{Jn} - \delta_n \frac{\sigma k}{\sigma-1} (\ln \tau^{np}_{Jn} - \ln \tau^p_{Jn}) \\
&- \frac{\sigma k}{\sigma-1} \ln d_{Jn} + \frac{k}{\sigma-1} \ln A_{Jn} + \ln \alpha_J - \frac{k-\sigma+1}{\sigma-1} \ln f_{Jn}.
\end{aligned}\tag{10}$$

Equation (10) illustrates how the removal of tariff uncertainty in destination n affects exports to that country. The terms in the first row of the equation's right-hand side denotes the log of expected tariffs. It is decomposed into two separate terms. The first term denotes the applied (preferential) tariff, and the second term measures the weighted log difference between potential and applied tariffs. In the following, this difference will be referred to as the tariff "gap" or the tariff "threat", exporters face in destination n. A removal of TPU implies that $\delta_n \to 0$, and that the tariff threat disappears. Exports to n increase, and since everything operates through a reduced ZPC, the adjustment takes place at the extensive margin.

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[5] $G(\varphi)$ suggests that the probability of drawing productivity $\varphi_j > \varphi_L$ decreases with the distance from φ_L. The parameter k raises the ratio to a positive number, $k > (\sigma-1)$, so that low-productivity firms and, hence, non-exporters are in the minority.

2.2 Separable Fixed Costs and Multilateral Trade

The baseline model offers no mechanism through which a removal of tariff uncertainty in another destination $m \neq n$ can have any impact on exports to n. To establish such a link, additional structure is imposed on the fixed market-entry costs, f_{Jn}. It is assumed that they can be separated into a local and a global component, $f_{Jn} \equiv f_n + f_J$.[6] With this assumption, the export decision of a firm is no longer independent from the total number of destinations it serves. The global fixed cost component, f_J, must be paid, irrespective of the number of destinations a firm serves. However, the burden can be distributed across sources of revenue.

Considering firm j's profits in all destinations $n = \{1, ..., N\}$ total export profits result as the sum of bilateral "partial" profits, $\tilde{\pi}$, minus *once* the global fixed cost:

$$
\begin{aligned}
\tilde{\pi}_{j1} - f_J &= (\tau_{J1}^E)^{-\sigma} d_{J1}^{-\sigma} \varphi_j^{\sigma-1} A_{J1} - f_1 - f_J \\
+\tilde{\pi}_{j2} &= (\tau_{J2}^E)^{-\sigma} d_{J2}^{-\sigma} \varphi_j^{\sigma-1} A_{J2} - f_2 \\
&\vdots \\
+\tilde{\pi}_{jN} &= (\tau_{JN}^E)^{-\sigma} d_{JN}^{-\sigma} \varphi_j^{\sigma-1} A_{JN} - f_N \\
\Pi_j(N) \equiv \sum_{n=1}^{N} \tilde{\pi}_{jn} - f_J &= \varphi_j^{\sigma-1} \sum_{n=1}^{N} \left[(\tau_{Jn}^E)^{-\sigma} d_{Jn}^{-\sigma} A_{Jn} \right] - \sum_{n=1}^{N} [f_n] - f_J.
\end{aligned}
\tag{11}
$$

The respective ZPC productivity Φ^* for exporting to all N destinations follows as

$$
\Phi_{JN}^* = \left[\frac{\sum_{n=1}^{N} [f_n] + f_J}{\sum_{n=1}^{N} (\tau_{Jn}^E)^{-\sigma} d_{Jn}^{-\sigma} A_{Jn}} \right]^{\frac{1}{\sigma-1}}.
\tag{12}
$$

2.2.1 General Implications of the Multilateral Productivity Threshold

Equation (12) states the productivity threshold required for exporting to all destinations. Whether this is optimal for a firm depends primarily on the partial bilateral profits.

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[6]This is similar to Hanson and Xiang (2011) who analyze the relative importance of global and local fixed costs. Focussing on US movie exports, they find that global fixed costs dominate. However, for manufacturing trade, they acknowledge that bilateral fixed costs must play a larger role, since trade patterns vary substantially across countries.

Lemma 2 *Irrespective of global fixed costs f_J, a firm j exports to a destination n only if bilateral partial profits are positive, $\tilde{\pi}_{jn} \geq 0$.*

This follows from Equation (11), where any row with negative partial profits lowers the total exporting profit. Hence, the number of destinations an exporting firm *potentially* serves results from an assessment of each market. To illustrate this, suppose that (11) ranks destinations in decreasing order of the bilateral partial profits, and note that this ranking is independent of the firm's individual productivity level. It then follows that

Lemma 3 *If $\tilde{\pi}_{j1} \geq \tilde{\pi}_{j2} \geq ... \geq \tilde{\pi}_{jN}$, and if global fixed costs can be covered, a firm exports to all destinations for which $\tilde{\pi}_{jn} \geq 0$ holds.*

Figure 1: Partial and Total Export Profits of Two Firms
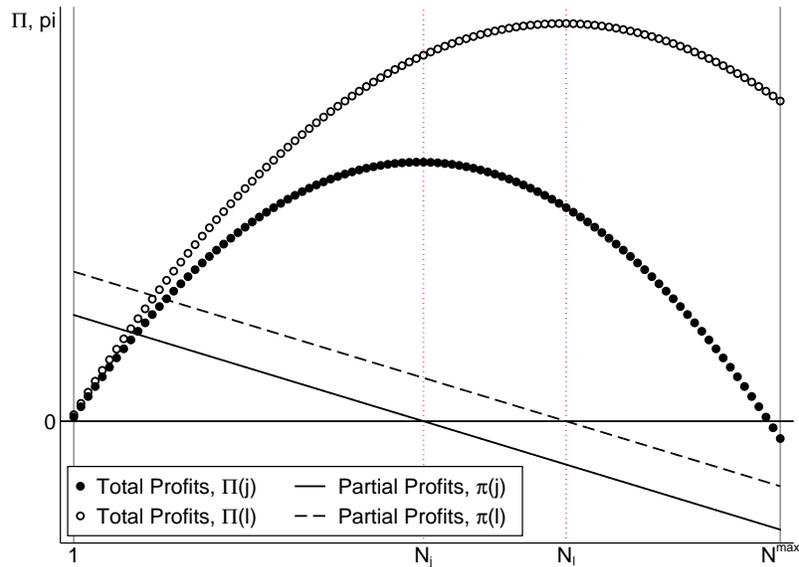
Note: Author's calculations based on Eq. (11) with 100 destinations.

Figure 1 summarizes Lemmas 2 and 3 considering two firms with different productivity levels. The horizontal axis denotes the range of potential export destinations n. They are ranked in decreasing order of the partial profits. The solid line with negative slope denotes these profits for firm $\varphi(j)$. The lower dotted curve denotes j's total export profits, conditional

on the number of destinations it serves. Total profits rise until partial profits become negative, which indicates its optimal number of destinations, N_j^*. The other firm, l, has productivity $\varphi_l > \varphi_j$ and serves a larger number of countries.

Lemma 4 *If $N = N^*$ denotes the optimal number of destinations to which a firm exports, then the productivity threshold Φ_{JN}^* is increasing in N.*

Lemmas 2-4 hold also for the case where fixed costs are purely country or product-destination specific. However, with a global fixed cost component it is possible that total exporting profits are negative in the first couple of destinations, and a firm will need the revenues from at least $N_j^{min} \leq N_j^*$ markets in order to export profitably. Tariff uncertainty in a large market may then prevent firms from exporting at all, because they expect to cover only a relatively small fraction of the global fixed cost burden by selling to it.

How this turns out shall be illustrated numerically, in the next subsection. Before doing so, the implications of this result should be briefly discussed. The prediction that firms start exporting only if they can serve a sufficient number of foreign markets seems to be at odds with market entry patterns of firms in high-income regions.[7] Nevertheless, things could be different for low-wage country exporters. Empirical evidence has repeatedly suggested within-product specialization across countries, where low-wage exporters tend to use different production technologies and factor proportions, yielding different "qualities" of a good (e.g. Schott, 2004; Khandelwal, 2010; Feenstra and Romalis, 2014). Theories on the organization of firms suggest that large-scale producers rely more on standardized production processes than firms offering customized products. The former are more exposed to the competitive pressures of globalization, because their activities can more easily be carried out in offshore locations (Thesmar and Thoenig, 2002; Holmes and Stevens, 2014). Along these lines, Amiti and Freund (2010) document that China's recent export growth was driven significantly by processing trade, where firms carry out production, but do not contribute to the development

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[7]For instance, Eaton et al. (2004, 2011) show that most French exporters serve only one foreign market, and that fewer and more productive firms export to more countries. The pattern suggests existence of product-destination specific fixed costs.

of a product or to distributing it to final customers. Manova and Zhang (2009) report patterns where private Chinese enterprises export more goods to more countries, while changing trade partners more frequently, than private foreign-invested firms. These patterns suggest that (at least some) Chinese exporters save destination-specific costs by focussing on standardized processes. The global fixed cost is paid to attract orders from firms that want to save labor costs. The bilateral fixed cost continues to be important, however, because it governs the eligibility a country to become part of an international production chain.

2.2.2 Bilateral Tariff Uncertainty and Multilateral Trade

Without further restrictions on how the global fixed cost component is allocated across sources of revenues, this subsection analyzes the implications of reduced tariff uncertainty in one destination. Number and characteristics of the destinations are modified in three steps.

Two Symmetric Countries. Supposing that firms consider exporting to two foreign destinations, $n = \{1, 2\}$, the baseline scenario describes the outcomes where tariffs are uncertain in country 1. While applied tariffs in country 1 correspond to the preferential rate $\tau_1^p = 1$, the non-preferential rate $\tau_1^{np} = 2$ might be applied with a probability $\delta_1 = 0.5$. Expected tariffs in country 1 are then $\tau_1^E = 1^{0.5} 2^{0.5} \approx 1.4$. There is no uncertainty in country 2 so that the expected tariff corresponds to the applied preferential rate $\tau_2^E = \tau_2^p = 1$. Equation (12) is used to compute ZPC productivity for exporting to country 1, to country 2, or to both countries. The elasticity of substitution is set to $\sigma = 3$, and, to obtain symmetry, all other non-tariff variables are set equal to 1.[8]

Columns (1)-(3) of Table 1 show ZPCs when firms export to both or only to one of the two countries, respectively. The first row shows the baseline scenario with tariff uncertainty in country 1. It suggests that, if any firm exports, it would most likely export to country 2, but not to country 1. Only firms with $\varphi \geq 3.53 \geq 3.18 = \Phi_2^*$ would find it profitable to serve

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[8] $\sigma = 3$ follows Handley and Limão (2013) who refer to estimates of Broda and Weinstein (2006). Other authors use $\sigma = 4$ (e.g. Head et al., 2014). As long as $\sigma > 1$ its absolute size has no impact on the qualitative results.

Table 1: Alternative Productivity Thresholds with Bilateral Tariff Uncertainty and Symmetric Countries

	(1) Φ_N^*	(2) Φ_1^*	(3) Φ_2^*
Baseline: $\tau_1^E = 1.4$	3.53	5.35	3.18
Treatment: $\tau_1^E = 1$	2.90	3.18	3.18

Note: Author's calculations based on Eq. (12) with two destinations, $\sigma = 3$, and all variables equal 1, except τ_1^E.

both destinations. Removal of tariff uncertainty, in the second row of the table, suggests that the least productive exporter would serve both destinations, because $\Phi_N^* = 2.90 < 3.18 = \Phi_2^*$. In fact, without tariff uncertainty, more firms are able to overcome the ZPC than in any case with uncertainty. Although policies in country 2 did not change, the removal of uncertainty in country 1 increases exports to that country.

Three Countries of Different Size. With three asymmetric countries $n = \{1, 2, 3\}$, it is possible to analyze how the size of the policy-making destination affects the multilateral threshold. Country 1 is again the one where tariffs are uncertain ($\tau_1^E = 1.4$). Three scenarios are considered where the size of country 1 is I. $E_1 = 1$; II. $E_1 = 2$; and III. $E_1 = 0.5$ in the respective cases. Country 2 is always large $E_2 = 2$ and country 3 is always small $E_3 = 0.5$. Besides this parametrization, and $\sigma = 3$, all other non-tariff variables equal one.

Figure 2 illustrates how the removal of tariff uncertainty translates to multilateral exports. It depicts the (log of the) Pareto-density function $g(\varphi)$. Most firms have a relatively low level of productivity and very few firms are very productive. The baseline scenario (with uncertainty in country 1) is normalized and shows the ZPC applicable for exporting to countries 1 and 2.[9] Only firms residing to the right of the baseline ZPC export under tariff uncertainty. The shaded areas indicate the amount of firms that become exporters when tariff uncertainty vanishes. It shows the smallest amount for case III, where country 1 is

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[9]The baseline ZPC for exporting to all three destinations is always higher but the mechanics remain the same.

Figure 2: Multilateral Productivity Thresholds with Bilateral Tariff Uncertainty and Asymmetric Countries

small. The largest amount of new exporters is found for case II, in which country 1 is large.

Although no allocation rule has been defined, the figure suggests that the size of a market is positively correlated with the fraction of global fixed costs it absorbs. If the policy-making country is very small, a removal of TPU has no effect at all. In fact, case III suggests that uncertainty in country 1 allows the least productive exporter to serve countries 2 and 3. When tariff uncertainty is removed the threshold for exporting to 1 and 2 equals the ZPC of exporting to 2 and 3. Hence, exports do not adjust at the multilateral level, because there is no additional entry of firms.

Market Size, Geography, and Fixed-cost Dominance. The final step feeds the model with actual data to assume the perspective of Chinese exporters. Market size, E_n, is now measured by the destination's share in world commodity imports (in billion US dollars). The country-specific fixed cost, f_n, is proxied by the distance between China's and the

destination's most populated cities (in 1,000 km), and the price level, P_n, is measured by the import price index.[10] The aim is to inspect which countries are most likely to exhibit a policy spillover, and to say something about the (relative) size of the global fixed cost.

Using data for 112 countries, ZPCs are computed for each destination, according to Equation (12), under the assumption of free trade (i.e., $\tau_{Jn}^E = d_{Jn} = 1$), and with $f_J = 0$. The results generate a ranking of countries' accessibility from China, based on its market size and geographic location. Figure 3 plots the computed Φ_ns against the destinations' share in Chinese commodity exports for the year 1999. The relationship is negative and almost log-linear. The estimated elasitcity of Φ with respect to $s_{n,99}^{CN}$ is about -0.37.[11] The EU-15 requires the lowest ZPC productivity, closely followed by Japan, South Korea, and the United States.

Figure 3: Computed Φ_n, and Destination's Share in Chinese Exports, 1999
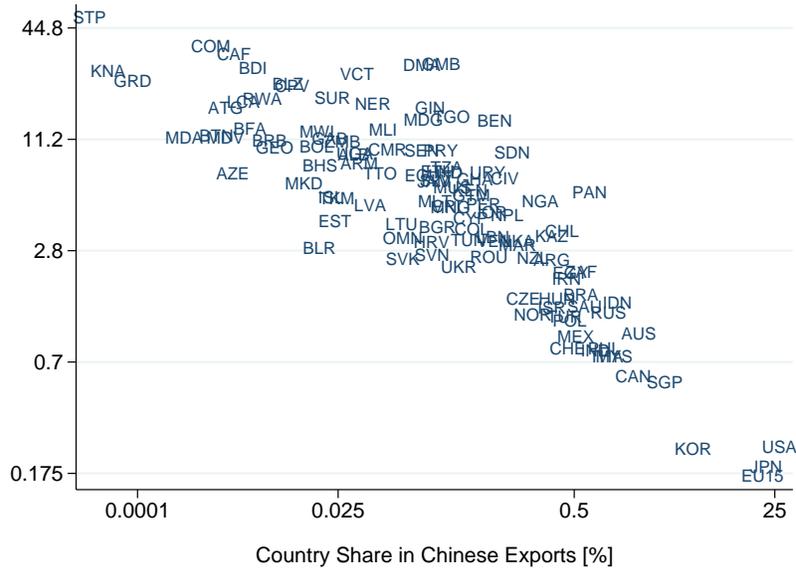
Note: Author's calculations, based on Equation (12), using data from UN Comtrade, CEPII GeoDist, and Penn World Tables 8.1.

With $f_J = 10$, which is close to the average $\bar{f}_n = \sum_{n=1}^N f_n/N = 9.68$, the least productive

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[10]The data is obtained from UN Comtrade, the CEPII GeoDist dataset, and Penn World Tables 8.1, respectively.

[11]Spearman's and Kendall's rank correlation tests reject independence of the computed ranking and the ranking of $s_{n,99}^{CN}$.

exporter could profitably sell abroad, if he accumulates the revenues from the first four destinations. The next best markets are too small or too distant to generate additional profits. Interestingly, the ranking is now different. The EU-15 remains first, but is now followed by the US, Japan, and South Korea. This suggests that market size governs the absorption of the global fixed cost.[12] In fact, generating rankings with increasing global fixed costs, $f_J = 0, ..., 100$, suggests that countries with large E_n become more accessible. A regression of computed Φ_n on f_J, and interactions of it with the other variables, yields the following elasticities:[13]

$$\ln \Phi_n = \underset{(0.11)}{1.16} + \underset{(0.02)}{2.86} \ln f_J - \underset{(0.00)}{0.13} (\ln f_J \times \ln E_n) + \underset{(0.00)}{0.02} (\ln f_J \times \ln f_n) - \underset{(0.01)}{0.25} (\ln f_J \times \ln P_n).$$

While the ZPC generally increases with f_J, the advantage of market size is indicated by the negative sign of the first interaction term. The advantage of geographic vicinity is considerably smaller, because it does not contribute to a destination's ability to cover the global fixed costs. High price levels increase the effective market size and therefore contribute to a lower elasticity of Φ_n with respect to f_J.

Returning to $f_J = 10$, and re-calculating the ZPCs under tariff uncertainty, with $\delta = 0.5$ and $\tau_{Jn}^{np} = 1.3$, indicates that TPU in one of the first four destinations leads to an increase of the lowest ZPC, but to no change when uncertainty prevails in countries $n = 5, ..., N$. Hence, if global fixed costs are about as high as the fixed cost of exporting to a destination 10,000 km away, the EU, the US, as well as Japan and South Korea, would be able to influence China's exports to third countries through its bilateral tariff policies. Using alternative values for f_J reveals that spillovers are possible even for $f_J = 3 < \bar{f}_n$, while $f_J = 100 > 10\bar{f}_n$ is not

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[12] A reason why the EU-15 continues to rank first is potentially the measurement error in the distance variable. It is likely to understate the actual distance from China to the European ports of entry, because it neglects the transportation mode and the distances of waterways. For example, the GeoDist dataset reports 10,993 kilometers between China and the US, and 7,831 kilometers between China and the Netherlands. Using data from http://www.sea-distances.org, instead, suggests 9,607 nautical miles between Zhanjiang and Rotterdam (via Suez Canal) but only 6,585 nautical miles to Los Angeles.

[13] Underset values denote standard errors of the coefficients. They are small, because Φ_n is a constructed variable.

sufficient to allow a fifth country to influence N^{min} for the least productive exporter.

Altogether, the numerical examples indicate that a spillover of bilateral policies on multilateral trade is possible, but that, given destination countries' characteristics in the actual data, only a small number of markets are sufficiently large to influence multilateral exports. Based on the results of this subsection, it will be assumed that the fraction of global fixed costs are divided and allocated to each destination, an exporting firm serves, and that the fraction, θ_n, allocated to a destination, is positively related with its effective market potential. This implies that a reduction of TPU in a large country $m \neq n$ increases its θ_m and decreases $\theta_{n \neq m}$. This redistribution induces entry of new firms into exporting, so that trade with both countries increases, even if policies in n remained unchanged. This effect should be observable at the extensive margin (i.e., a good is exported to more destinations). Moreover, if the selection of firms does not entail any dynamic gains, the effect should phase out after a transitional period of adjustment.

3   Empirical Application

This section describes how the predictions of a policy spillover are empirically evaluated. Following previous studies on the bilateral effects of TPU reductions, the removal of US tariff uncertainty in conjunction with China's WTO accession is exploited to analyze its effect on China's exports to the EU.

3.1   China-US and China-EU Trade Relations

The trade policy environments China faced in US and European markets allow for an analysis of a US policy spillover. The US and China established their diplomatic relations in the late 1970s, when Deng Xiaoping initiated economic reforms. In 1980 the US granted China preliminary MFN status for its exports, which implied a reduction of the tariffs applied before in corresponence to the "Column 2" schedule the US typically applies to non-market

economies. Column 2 rates were originally defined under the Smoot-Hawley Tariff Act of 1930, and higher than the MFN rates, which were gradually dismantled during GATT/WTO negotiations (Pierce and Schott, 2013). The preliminary nature of China's MFN status, however, entailed the risk that the US would return to apply Column 2 rates, if the US Congress would vote against its continuation. Hence, MFN rates were guaranteed for one year, but uncertain for those that would follow.

Handley and Limão (2013) quote a number of business practitioners and politicians, supporting the view that TPU had a deterring effect on investment into Chinese exports to the US. In the 1990s China's MFN status was actually close to being withdrawn. After the Tiananmen Square incident, in 1989, political opposition to China's MFN status arose due to concerns about the violation of human rights standards. Pierce and Schott (2013) report that, at this time, votes sufficed for a return to Column 2 tariffs, but the US Senate failed to act on this. Shortly before China's WTO entry, in the years 1997-2001, the votes against its MFN status still amounted to 38 percent, on average.[14] Throughout the 1990s, and beyond, political tensions between the two countries remained: In 1999, NATO accidentally bombed the Chinese embassy in Serbia, and, in 2001, China refused to return a US surveillance plane after a collision with a Chinese fighter jet over the South China Sea.

While diplomatic tensions appear to persist until today,[15] the preliminary status of China's MFN treatment was turned into "permanent normal trade relations" (PNTR), upon its accession to the WTO in December 2001. This removed the inhibiting effect of TPU for Chinese exporters and encouraged their entry into the US market (Handley and Limão, 2013; Feng et al., 2014). The fact that the preliminary MFN status was never *actually* overturned makes this policy change appropriate for a natural experiment.

The policy environment in Europe has been, from the viewpoint of Chinese exporters, relatively stable. China and the former European Community (EC) agreed on an equivalent to PNTR in 1979, which established China's MFN status in Europe. In addition, China

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[14]At least 50 percent were required to overturn China's MFN status.

[15]Most recently, the US destroyer USS-Lassen crossed territory in the South China Sea, claimed by China.

became a beneficiary country under the Generalized Scheme of Preferences (GSP) in 1980. The GSP grants preferential market access to developing countries through discounts on applied MFN rates. In contrast to the China-US relations, the European GSP entails a lower degree of tariff uncertainty, as it sets out tariff preferences for several years. Although the possibility exists that GSP rates are withdrawn, and replaced by MFN tariffs, it is presumably more predictable. Graduation from GSP rates happens when a country has reached a certain level of economic development, or when it has become a dominant exporter of a good so that it threatens market access for other GSP beneficiary countries. Most importantly, however, is the notion that nothing changed in these formal procedures when China entered the WTO.[16]

3.2 Empirical Framework

To apply the theoretical predictions to the event China's WTO accession and the simultaneous US policy change, the empirical specification adopts a difference-in-difference (DID) strategy. The estimation equation is derived from the previous section, which extends the bilateral gravity equation, stated in (10), by the fraction of global fixed costs allocated to destination n:

$$
\begin{aligned}
\ln R_{Jn} = \quad & -\frac{\sigma k}{\sigma - 1} \ln \tau_{Jn}^p - \delta_n \frac{\sigma k}{\sigma - 1} (\ln \tau_{Jn}^{np} - \ln \tau_{Jn}^p) \\
& -\frac{\sigma k}{\sigma - 1} \ln d_{Jn} + \frac{k}{\sigma - 1} \ln A_{Jn} + \ln \alpha_J - \frac{k - \sigma + 1}{\sigma - 1} \ln (f_n + \theta_{Jn} f_J). \quad (13)
\end{aligned}
$$

The parameter $0 < \theta_{Jn} \leq 1$ indicates that a part of f_J is allocated to each destination n, Chinese firms serve. $\theta_{Jn} = 0$, if China does not export this good to n. The US policy change affects θ_{Jn}. Reduction of TPU in the US induces $\delta_{US} \to 0$, so that $\theta_{J,US}$ increases and θ_{Jn} decreases for $n \neq US$. From $\partial R_{Jn}/\partial \theta_{Jn} < 0$ it follows that exports to country n increase. As long as there is no tariff uncerainty in country n itself ($\delta_n = 0$), the second term on the right-hand side of the equation disappears, and only applied tariffs exert their negative effect

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[16]China's GSP status in the EU was withdrawn in 2014, when it had been ranked as a higher-middle income country for three consequtive years.

on exports.

The estimation equation summarizes several variables, using appropriate fixed effects. This is convenient, because some variables the model suggests might not be measurable in an appropriate way, while others that should be included, might not have been identified by the theory. Accordingly, equation (13) is rewritten in product, destination, and time dimension:

$$\ln R_{Jnt} = b_1(GAP_J^{US} \times D_t^T) + b_2 \ln \tau_{Jnt} + b_{Jn} + b_{nt} + b_{St} + \varepsilon_{Jnt}. \tag{14}$$

The first term in this equation denotes the US policy change, which is assumed to reduce the size of θ_{Jn}. The variable GAP_J^{US} measures the US tariff threat, specific to product J, which was perceived by Chinese exporters prior to WTO accession. The variable D_t^T is a WTO dummy, which takes values of zero in the years $t \leq 2001$ and one for $t > 2001$. As b_1 quantifies the effect of the *removal* of US tariff uncertainty, it should be positively different from zero, if a policy spillover exists. Moreover, if b_1 measures correctly the elasticity of R_{Jn} with respect to $(f_n + \theta_{Jn}f_J)$, it corresponds to $(k - \sigma + 1)/(\sigma - 1)$. Using estimates of $k = 4.854$ and $\sigma = 4$ from Head et al. (2014) gives $b_1 \approx 0.618$, as a reference value.

The other measured variable, τ_{Jnt} denotes the applied tariff rate. In the EU it varies across products and over time. Its estimated coefficient, $\hat{b}_2$, should also be negative. The remaining terms b_{Jn}, b_{nt}, and b_{St} denote product-destination, destination-time, and sector-time fixed effects, repectively. They absorb the time-invariant ad-valorem transportation costs and preference structures in country n, time-varying aggregate conditions in each destination, and sector-specific time-varying intercepts. The aggregation of the latter is necessary to identify the impact of the US policy change, which varies in the same dimensions as α_J. Several robustness checks will be applied to provide support for a distinct effect of GAP_J^{US} on Chinese exports to $n \neq US$.

3.3   Data and Measures

Estimating equation (14) requires information on US MFN and Column 2 tariffs, before China's WTO entry, to construct GAP_J^{US}. The data is drawn from Feenstra et al. (2002) and reports ad-valorem equivalents of applied tariffs at the 8-digit level of the Harmonized Tariff Schedule (HTS). Aggregation to the 6-digit level allows matching the data with product codes of the Harmonized System nomenclature (HS6). Each HS6 product category will correspond to a product J. The US tariff threat is calculated at the 8-digit level, as the log difference between Column 2 and MFN rate, and is subsequently aggregated to the HS6 level. Tariffs are coded as $\tau = 1 + (\%\text{-rate}/100)$. Following Pierce and Schott (2013), the year 1999 is selected as benchmark:

$$GAP_J^{US} \equiv \frac{1}{H(J)} \sum_{hts=1}^{H(J)} \ln \tau_{hts,99}^{Col2} - \ln \tau_{hts,99}^{MFN},$$

where $H(J)$ denotes the number of HTS products comprised in the respective HS6 product category J.[17]

Figure 4 depicts the variation of GAP_J^{US} within sectors, S, which are comprised of the HS Sections IV through XX. In each of these sectors, the median product has about $GAP^{US} \geq 0.2$, and in all sectors but apparel and footwear there is at least one good with $GAP^{US} = 0$. This relatively equal distribution of tariff gaps across sectors is reassuring against concerns that the US policy change affected only one single industry.[18]

Destination countries, n, reflect each of the EU-15 member states, as of 1995. Information on applied EU tariffs on Chinese products is obtained from the TRAINS tariff schedules, accessible through the World Integrated Trade Solution (WITS), which report ad-valorem equivalents at the 8-digit level of the Combined Nomenclature (CN8). At the 6th digit they

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[17]All product codes of the HS6 nomenclature are converted into the codes appearing in the 1988/1992 HS classification. The respective correspondence tables were accessed through the United Nations Statistics Division (UNSD).

[18]As a standard extension to the baseline equation (14), regressions will include information on textiles, clothing, and apparel products, which benefited from removal of import quotas in the EU.

Figure 4: GAP_J^{US}, across HS6 products within Sectors, 1999
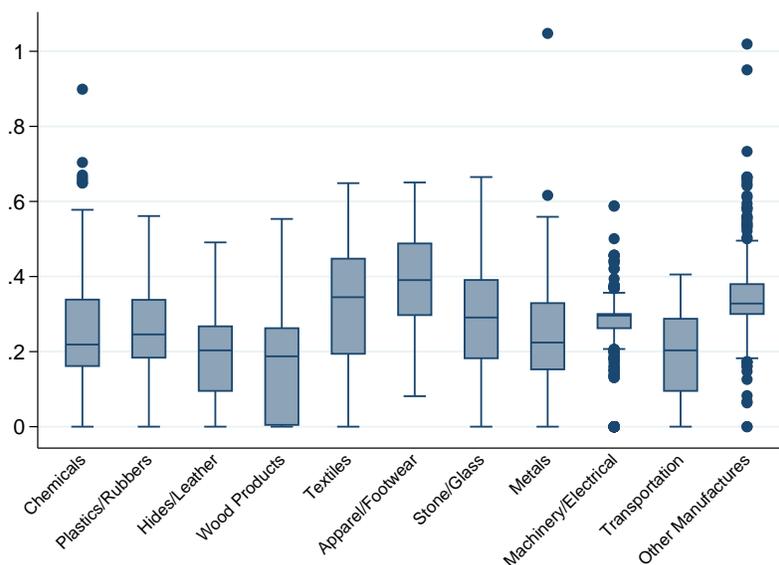
Note: Author's calculations based on US Tariff Data (Feenstra et al., 2002). Horizontal axis denotes manufacturing sectors, according to HS Sections IV through XX.

can be matched with HS6 products J. A challenge with this data is that information on China- and product-specific exemptions from the GSP are published only every couple of years, in the repective European Commission Regulations. Correspondence tables provided by Bernard et al. (2012), and more recent versions from the EUROSTAT Reference and Management of Nomenclatures (RAMON) archive, are used to trace such goods over time.[19] With the complete set of applied and applicable tariffs it is possible to compute the tariffs faced by China even when it does not export to the EU. This information is required to assess the entry of Chinese exporters into European markets.

The dependent variable, R_{Jnt} uses data on European countries' imports from China, accessed through the UN Comtrade Online Database. Due to changing reporting practices around China's WTO entry, Belgium and Luxembourg are treated as one country. To distinguish the intensive from the extensive margin of trade, missing information on imports from China, for a given HS6-destination-year combination, is assumed to reflect that no trade had

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[19]Details on this procedure, and on the construction of all other variables are provided in Appendix B.

taken place.

Table 2: China's Manufacturing Exports to the EU: Tariffs, Values, and Growth; 1997-2009

		Cross-sectional averages			Within product-destinations	
Years	WTO	Tariffs	Trade	Exports	Δ Exports	Δ Tariffs
1997-2001	No	3.27	0.47	1.95	0.19	-0.21
2002-2006	Yes	2.98	0.61	3.86	0.32	0.30
2007-2009	Yes	2.91	0.73	7.38	0.15	-0.43

Note: Author's calculations based on UN Comtrade, WITS, and EC Regulations. WTO denotes China's membership status; Tariffs denote average percentage rates in existing trade relationships, Exports measured in million US dollars (current prices). Growth rates denote average normalized rate (Davis et al., 1998, see explanation in the next section).

Table 2 presents summary statistics for values (in current million US dollars) and tariffs (in percent) of China's EU exports, in three periods. The first data column shows how average applied tariffs across existing trade relationships decreased only very slowly, by 0.29 percentage points between 1997-2001 and 2002-2006, and by another 0.07 percentage points in the last period. Meanwhile, the fraction of active (non-zero) trade relationships increased from 0.47 to 0.73, within little more than a decade, whereas the average export value almost doubled in the first four years since WTO entry. Within product-destination pairs, Jn, the last two columns show how export growth accelerated during the years 2002-2006, although tariffs for these trade relationships had been increasing, which presumably reflects graduation from GSP tariffs. In the final dataset, the average product J is associated with a US tariff threat of about 0.272 log points. 216 of the 3,271 HS6 products (i.e., 5.80 percent) reveal to have $GAP_J^{US} = 0$. The tariff threat for an average *exposed* product is 0.288.

4 Results

This section presents the results of extensive analyses regarding the existence, robustness, and further characteristics of the US policy spillover on China's EU exports.

4.1 Main Findings

4.1.1 Level of Chinese Exports to the EU

Table 3 presents results from estimating Equation (14) for the period 1995-2005. The baseline sample considers the full range of manufacturing products of which there are $3,271$, comprised in the HS Chapters 28-96.

Column (1) reports a positive and statistically significant coefficient for the removal of US tariff uncertainty. It suggests that Chinese exports of threatened products increased by 19.4 percent relative to a non-threatened product. Columns (2) and (3) take into account potentially differential patterns in the textile and clothing sector (T&C), where the EU removed import quotas for Chinese goods in 2002, 2005, and 2009 (Utar, 2014). These goods are comprised in HS Chapters 50-67 and were excluded in the estimation reported in column (2), whereas column (3) includes dummy variables for the products J subject to quotas removals in the respective years.[20] The estimated effects are somewhat lower but still statistically significant and positive. It suggests that China's exports increased due to both the removal of quotas and the removal of US tariff uncertainty. The size of the estimated coefficients, $\hat{b}_1$, range around the reference value of $b_1 \approx 0.618$ suggested in the previous section.

The remaining columns, (4) through (6), replace the continuous measure GAP_J with a discrete variable, similar to Handley and Limão (2013). GAP is divided into four groups. The first group considers the goods where $GAP = 0$, which is defined as the base value. The second group considers the bottom quartile, where $GAP = [0.001, 0.209]$, the last group includes the top quartile with $GAP = [0.356, 1.019]$, while the intermediate group summarizes values ranging from 0.209 to 0.356 and comprises about half of the products. The results of the baseline specification are qualitatively confirmed, while the coefficients suggest a slight non-linearity. This is similar to the findings of Handley and Limão (2013) for China's US

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[20]Information on the affected products is available online at the Système Integré de Gestion des Licences à l'Exportation et à l'Importation (SIGL). See Appendix B.

Table 3: Chinese Exports to the EU-15 and the Removal of US Tariff Uncertainty; Linear and non-parametric panel estimation, 1995-2005

	Baseline			Discrete Measure		
	(1)	(2)	(3)	(4)	(5)	(6)
Product Range:	Full	No T&C	Full	Full	No T&C	Full
US Tariff Threat	0.674**	0.437**	0.660**			
	(0.066)	(0.079)	(0.066)			
$p[> 0]$-$p[25]$				0.304**	0.292*	0.306**
				(0.067)	(0.071)	(0.067)
$p[25]$-$p[75]$				0.457**	0.444**	0.456**
				(0.065)	(0.069)	(0.065)
$p[75]$-$p[100]$				0.518**	0.422**	0.514**
				(0.066)	(0.071)	(0.066)
EU Tariff	-1.526**	-1.271**	-1.481**	-1.516**	-1.252**	-1.470**
	(0.266)	(0.275)	(0.266)	(0.266)	(0.275)	(0.266)
EU Quota '02			0.107**			0.108**
			(0.038)			(0.038)
EU Quota '05			0.183**			0.180**
			(0.044)			(0.044)
Observations	290,761	222,385	290,761	290,705	222,329	290,705
R-squared	0.233	0.241	0.233	0.233	0.242	0.234

Note: Table shows estimates based on Eq. (14), and data from various sources (see text). Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year, included in all regressions. Robust standard errors in parentheses; significance: [a] $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

exports. The remaining variables produce the expected signs throughout all specifications. The coefficient for import tariffs, however, seems to be somewhat too small. A potential explanation is its little variation in many product categories, where tariffs had been low even at the beginning of the sample period.

4.1.2 Extensive vs. Intensive Margin

This subsection analyzes whether the US policy change has facilitated the creation of new trade relations between China and EU countries. Table 4 presents results of different specifications that capture adjustments at the extensive product-destination margin. In contrast to the baseline estimation, tariffs are included with a one-period lag, i.e. τ_{Jt-1}.

The first two columns present the odds-ratio and the coefficient of a logit model, where the

Table 4: Chinese Market Entry in the EU-15 and Removal of US Tariff Uncertainty; Alternative Specifications, 1995-2005

| | Logistic Regressions | | Linear Regressions | | |
| | Odd Ratio | Coeff. | # Destinations | Norm. Growth | Log Growth |
	(1)	(2)	(3)	(4)	(5)
US Tariff Threat	3.399**	1.223**	0.915**	0.086**	-0.040
	(0.299)	(0.088)	(0.243)	(0.030)	(0.027)
EU Tariff	0.417*	-0.875*	-1.337^a	-0.934**	0.047
	(0.169)	(0.404)	(0.766)	(0.246)	(0.191)
EU Quota '02	1.155*	0.438**	0.635**	0.134**	0.120**
	(0.103)	(0.066)	(0.165)	(0.018)	(0.015)
EU Quota '05	1.485**	0.396**	0.151	0.398**	0.388**
	(0.120)	(0.081)	(0.191)	(0.035)	(0.030)
Observations	265,930		37,480	301,789	230,134
R-squared	0.216		0.442	0.017	0.017
Fixed effects	Jn, t		J, St	Jn, nt, St	Jn, nt, St

Note: Table shows estimates based on alternative specifications. Columns (1) and (2) show odds ratio and coefficients of logit estimates. Column (3) considers the number of EU-15 destinations served by China. The last two columns compare the normalized growth rates, conditional in inclusion (4) and exclusion of the extensive margin (5). Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year. Standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

dependent variable equals one whenever China exports to a given product-destination pair, and zero otherwise. Column (1) suggests that the *removal* of trade barriers (i.e., US tariff uncertainty and EU import quotas) increases the probability of observing Chinese exports. Higher tariffs make trade less likely to occur. The coefficients stated in column (2) show the marginal effects. Tariffs reduce the probability of exporting to the EU, whereas the removal of other trade barriers reveals positive and statistically significant effects. Column (3) considers an alternative specification where the dependent variable measures the number of EU destinations to which China exports good J. This reduces the dataset to the product-time dimension. The results confirm that products with a high US tariff threat were exported to more destinations after uncertainty was removed. Columns (4) and (5) consider a third test against adjustments at the extensive margin. They report the estimated effect on the annual growth rate of Chinese exports in a particular product-destination pair. Instead of computing the growth rate as the log-difference, which is unable to capture entry and exit,

the normalized growth rate is used (Davis et al., 1998; Pierce and Schott, 2013).[21] In column (4) the dependent variable corresponds to this growth rate, whereas column (5) removes observations where it takes the bound values denoting entry and exit. The results suggests that an effect of the US policy change cannot be confirmend when the extensive margin is neglected. Together, the main findings cannot reject the existence of an impact of US policies on China's multilateral export performance.

4.2 Robustness

This subsection extends the baseline specification by various controls to see whether important variables have been neglected, which could have led to a false rejection of the null hypothesis that $b_1 = 0$. The results are summarized in Table A1, in Appendix A. The Appendix also provides information on the correlation among control variables (Table A2).

4.2.1 Unobserved Productivity Shocks

The first set of control variables accounts for increased investments in product categories, where China has revealed a comparative advantage over the past decades. The basis of this measure is its revealed comparative advantage (RCA), as reported in the CEPII RCA dataset for the years 1995-1997 and 2008-2010. The ratio, RCA_{08-10}/RCA_{95-97} constitutes a product specific measure of structural change which is interacted with the WTO dummy. Columns (1) and (2) of Table A1 indicate that the estimated impact of the US policy change remains highly significant with $\hat{b}_1 = 0.503$, while the RCA change produces a complementary positive coefficient. Interacting the RCA-ratio with the exchange rate of China's currency, in terms of the European countries' respective currency units, suggests further that depreciations of the Yuan have not contributed to confounding effects regarding the US policy change.

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[21]The rate is defined as

$$g^N \equiv \frac{R_t - R_{t-1}}{0.5(R_t + R_{t-1})},$$

and takes the bound values, $g^N = \{-2, 2\}$, upon exit and entry, respectively.

The next two columns use a more direct measure of productivity shocks: quality-adjusted export prices (Feenstra and Romalis, 2014). Investments into better technologies and other productivity shocks should be reflected by lower average prices of an export good, which then boosts exports to other countries. Computing the ratio of these prices from the years before and after China's WTO accession (1997-1999 vs. 2002-2004), and interacting it with China's WTO dummy, produces the results shown in column (3). It does not appear that the rejection of $b_1 = 0$ is driven by unobserved productivity shocks which are reflected in lower prices. Column (4) discounts these export price changes by the destinations' respective import price changes and confirms the previous result.

4.2.2 Chinese Policies

In conjunction with its WTO accession, also China complied to policy changes which may have had a positive effect on its multilateral export performance. Three aspects shall be taken into account: (i) the deregulation of foreign investment restrictions; (ii) dismantlement of import tariffs; and (iii) export subsidies.

To investigate the changes in foreign investment restrictions, the analysis follows Pierce and Schott (2013) by evaluating exports of products with a relatively high contracting intensity. This measure was developed by Nunn (2007) and denotes the fraction of production inputs which are defined as differentiated goods in the classification of Rauch (1999). Column (5) is uses the strict measure, where only the fraction differentiated goods is measured. Column (6) uses a broader measure, which takes into account also less differentiated (reference-priced) inputs. In both cases, a positive impact of China's foreign investment deregulation can be confirmed, but they do not appear to influence the estimated effect of the US policy change.

Tariff reductions in China have been shown to benefit exporters who switched to sourcing production inputs from abroad (Feng et al., 2012). To account for such changes in the present paper, Chinese tariff changes in HS6 product categories are mapped to their associated

output goods, J, using "Make and Use" tables of the US economy for the year 2002. The procedure is similar to the generation of the contracting intensity measures (Nunn, 2007), and is described in Appendix B. The resulting measure reflects a weighted average of changes in average tariffs on J's inputs between 1996-2000 and 2001-2005. Likewise, weighted averages of changes average imports of the production inputs are computed. The results are shown in columns (7) and (8), respectively, and suggest that tariff reductions have indeed contributed to China's export performance. Imported inputs, however, do not appear to be statistically significant. The effect of the US policy change remains in a comparable order of magnitude as in the previous specifications.

Although export subsidies by the Chinese government are not directly related to the date of its WTO accession, data on anti-dumping (AD) investigations and countervailing duties (CVD) Bown (2014, 2015), initiated by European firms, may shed light on their relevance in determining China's export boom in the early 2000s. Each product subject to approved and ongoing AD investigations, after the year 2003, and those subject to CVD are interacted with China's WTO dummy to inspect whether their exports have shown a differential performance since 2002. The results in column (9) confirm such an effect for AD investigations but only little evidence for CVD, which is reported only for the years 2010 and later. The estimate for the US policy change remains quantitatively unaffected.

4.2.3   European Events

Focussing on European events that were not related but occurred simultaneously to China's WTO accession, potential crowding-out effects of the European currency union are investigated. Southern European economies (i.e., Greece, Italy, Spain, and Portugal; GIPS) experienced a decline in their inflation rates and in the interest rates on government bonds, as they switched from their domestic currencies to the Euro. While cheaper borrowing may have contributed to misallocation of capital in these countries (Feldstein, 2012; Gopinath et al., 2015), rent-seeking investors were incentivized to move capital to other countries. Moreover,

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past investments in the GIPS countries, denominated in their currencies, faced higher than expected debt services. Hence, Chinese exporters may have benefitted from the EURO introduction through subsitution effects of import sources in some European countries (e.g. Dauth et al., 2014). To evaluate this possibility GIPS countries' RCA for the years 1999-2001 is interacted with the WTO dummy. The results in column (10) confirm such an effect, but $\hat{b}_1$ remains positively different from zero.

The next test suggests that the effect of the US policy change might materialze in only one or a few countries of the EU. For instance, the European integration process, with its removal of tariffs and other trade barriers towards countries entering the EU during the enlargement rounds since 2004, could have boosted Chinese exports to those countries, while shipping their goods to EU-15 destinations in the first place. To this possibility, equation (14) is repeatedly estimated while removing each time one of the $N = 14$ destination countries. Figure A1 shows the estimated coefficients and confidence intervals when the respective country on the vertical axis is removed from the sample. There is no statistically significant variation in the coefficients.

Finally, to see whether any other Europe-specific event has driven the rejection of $b_1 = 0$, the estimation equation is applied to data on Chinese exports to Japan. Unlike to the EU, Japan had not imposed any import quotas on Chinese textiles, clothing, or apparel products, so that only the applied tariffs must be taken into account. The results in column (11) of Table A1 indicate that the US policy change had a positive and statistically significant impact on China's exports to Japan. The estimated coefficients of both the tariff threat and applied tariffs are in similar orders of magnitude as in the EU case.

Altogether, the robustness checks reveal that the US policy change was not the only event that has contributed to China's export boom following its WTO accession. Nevertheless, the results of the baseline specification appear to be statistically and quantitatively robust. This is also supported by the results in the last column of Table A1, where several control variables are simultaneously included.

4.3 Further Results

This subsection provides further insights on the characteristics of the effect the US policy change had on China's EU exports. It first investigates the evolution of the policy spillover over time, using an extendend sample period. The second step addresses the assumption that exports to a country n increase when the fraction of fixed costs allocated to another (large) destination, $\theta_{m \neq n}$, rises. Finally, the possibility of heterogeneous elasticities of exports with respect to the US policy change is investigated.

4.3.1 Dynamic Adjustments and Transitional Growth

Available trade and tariff data allows extending the analysis up to the year 2012, so that the effect can be studied over time. Re-estimating Equation (14) for the years 1995 up to 2012, using GAP_J^{US} and taking into account the removal of import quotas, produces a coefficient $\hat{b}_1 = 0.748$ (see Table A3). It is larger and statistically different from that obtained for the baseline period shown in Column (3) of Table 3. One explanation could be that the US policy change interacts with a dynamic component. The dashed line in Figure 5(a) depicts the evolution of $\hat{b}_1$ as the sample period increases.

A dynamic specification should eliminate the positive correlation between estimates of b_1 and the length of the post WTO-entry period. However, including the lagged dependent variable $\ln R_{Jnt-1}$ on the right-hand side of the estimation equation creates problems, due to correlated errors which lead to biased coefficients. Following Roodman (2009), this bias is controlled by inferring the upper and lower bounds of the true coefficient of $\ln R_{Jnt-1}$, which should lie between the estimate obtained from a dynamic pooled OLS model (POLS) and a dynamic fixed effects (FE) model. The former overestimates the true coefficient, while the latter produces a downward biased estimate for the lagged endogenous variable. If $\hat{b}_1$ is positively correlated with the dynamic component, a downward-biased dynamic specification will pick up some but not all of the dynamics that were induced by the US policy change.

The left panel of Table A3 confirms this expectation. It shows the baseline in column (1),

the dynamic FE in column (2), and the dynamic POLS estimate in column (3). In the two dynamic specifications, the removal of the US tariff threat reveals a lower coefficient for GAP than in the baseline, but the POLS model produces unplausible results for the effect of EU tariffs. It seems that the FE model in column (2) is more plausible. The solid line in Figure 5(a) shows the point estimates of b_1 in the dynamic FE specification. Its evolution suggests that initial estimates are corrected downwards as observations in the post-WTO period effect increase. For period lengths of 1995-2006, and beyond, the estimates converge to a level effect of $\hat{b}_1 \approx 0.20$. With these estimates in hand, the long-run effect of the US policy change can be computed. If γ denotes the coefficient for the lagged endogenous variable, $\hat{b}_1^{long} = \hat{b}_1/(1-\gamma)$ can be compared to the benchmark from Head et al. (2014). Column (2) of Table A3 suggests $\hat{b}_1^{long} = 0.380 < 0.618$, indicating that the US policy change does not perfectly measure the change in the fraction of global fixed costs allocated to the EU.[22]

Figure 5: Evolution of the US Policy Effect on Chinese Exports to the EU; 2002-2012

(a) Level Effect   (b) Growth Effect
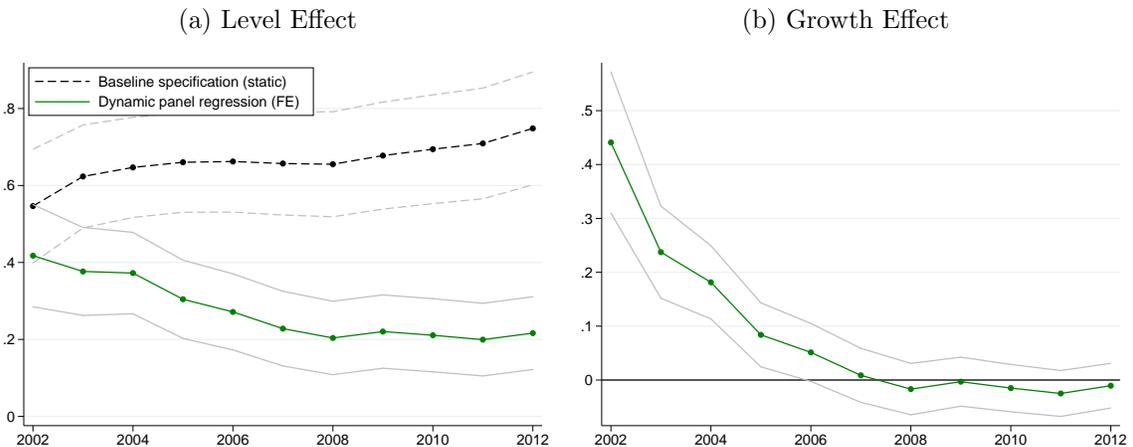
Note: Figure shows estimated effects of GAP_J^{US} on the level (Panel a) and normalized growth rate (Panel b) of Chinese exports to the EU-15, as obtained from varying period length 1995-T ($T = 2002, ..., 2012$). All regressions include product-country, sector-year, and country-year effects.

Turning to the right panel of Figure 5, the evolution of the US policy effect on China's export growth can be inferred. Since the theoretical exposition in section 2 is static, and

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[22]An explanation could be that the reference value for b_1 reflects the elasticity of R_{Jn} with respect to $\ln(f_n + \theta_{Jn} f_J)$, while GAP_J^{US} is expected to affect only the product-specific fixed cost. Thus, the reference value should be seen as an upper bound for the estimated b_1.

suggests that less productive exporter are selcted into foreign markets, an enduring growth effect must not be expected. In fact, Figure 5(b) indicates that the US policy change induced a transitional growth trajectory, which phases out after a few years. This is in line with the view that the reduction of US tariff uncertainty has benefitted exporters which focus on standardized production with high fixed costs, but little on product development.

4.3.2   US Share in Chinese Exports

Investigating further the theoretical assumptions that motivated the analysis of the policy spillover, this subsection focuses on the mechanism of allocating global fixed costs f_J across destinations. The main findings were based on the direct inclusion of GAP_J^{US} into the regression equation, while the presumed change in $\theta_{J,US}$ has been neglected. To provide (suggestive) evidence for a fixed-cost reallocation, $\Delta\theta_{J,US}$ will be proxied with an alternative measure. This is done by computing the share of China's exports of product J shipped to the United States, s_{Jt}^{US}, in the periods before and after China's WTO entry. The pre-entry period is denoted as $t = 1992, ..., 1996$ whereas the post-entry period covers the years 2004-2008. The difference, $\Delta\bar{s}_J^{US} = \bar{s}_{J,US}^{post} - \bar{s}_{J,US}^{pre}$, serves as the proxy for $\Delta\theta_{J,US}$, and is used to replace GAP_J^{US} in equation (14). A positive coefficient lends support to the hypothesis that global fixed costs are redistributed across destinations, according to their absorptive capacities.

Table 5 shows the results obtained from a sample period 1995-2005 which corresponds to the baseline. Column (1) is analogous to the third column of Table 3 while the remaining columns can be compared with Table 4. The estimated coefficients for EU tariffs and the quota removals in the first column are qualitatively and quantitatively unchanged. The coefficient obtained for the US share in Chinese exports is positive and statistically significant, but lower than that obtained from the original specification. This downward bias may be attributed to measurement error in the proxy for $\Delta\theta_{J,US}$. Columns (2) through (6) confirm the qualitative results from the main findings. Together, the results are reassuring that Chinese exporters benefit from selling much to the US, as this reduces the *de facto* fixed cost

Table 5: Chinese Exports to the EU-15 and the Share of Exports shipped to the US; Alternative Estimators, 1995-2005

	Baseline	Logit		Linear Regressions		
	Levels	Odd Ratio	Coeff.	# Dest.	Norm. vs.	Log Growth
	(1)	(2)	(3)	(4)	(5)	(6)
US Share	0.274**	1.233*	0.210*	0.665**	0.059**	0.035^a
	(0.050)	(0.112)	(0.091)	(0.191)	(0.022)	(0.021)
EU Tariff	-1.406**	1.336	0.290	-1.611*	-0.932**	0.018
	(0.266)	(0.753)	(0.563)	(0.722)	(0.247)	(0.192)
EU Quota '02	0.128**	1.845**	0.612**	0.657**	0.136**	0.117**
	(0.037)	(0.168)	(0.091)	(0.165)	(0.018)	(0.015)
EU Quota '05	0.188**	1.826**	0.602**	0.149	0.398**	0.387**
	(0.044)	(0.340)	(0.186)	(0.193)	(0.035)	(0.030)
Observations	290,126	154,780		36,880	298,131	227,275
R-squared	0.233	0.348		0.445	0.071	0.071
Fixed effects	Jn, nt, St	Jn, t		J, St	Jn, nt, St	Jn, nt, St

Note: Table shows estimates based on Eq. (14), but with a different treatment variable instead of GAP. Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year. Robust standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

burden for exporting to European markets.

As alternative specifications, changes in the share of product J exports shipped to other countries are used. Table A4 in Appendix A presents the estimated effects a reallocation of global fixed costs to Canada, Japan, New Zealand, and South Korea has on Chinese exports to the EU. Increased allocation to small and distant economies, such as New Zealand in column (3), do not appear to have an impact on exports to the EU. This is in line with the theoretical prediction that only large countries are able to exhibit spillovers. The effect for allocation to Canada, shown in column (1), is positive and statistically significant at the 10 percent level. This suggests that its market is just large enough to absorb a sufficient fraction of f_J that facilitates exports to third countries. However, tight economic relations between Canada and the US might be driving this effect. In contrast, column (2) and (4) suggest that increasing concentration of exports to relatively nearby destinations in Japan and South Korea has an inhibiting effect of exports to the EU. This suggests that the bilateral fixed costs f_n are more important for Chinese exports to these countries, which would imply that the goods

traded with Japan and South Korea are different from thos exported to the US, Canada, and the EU. One reason for such a segmentation could be the frequent trade within production networks among East-Asian countries, while Chinese exports to distant high income markets focus on finished goods. The next subsection provides suggestive evidence that goods differ in their likeliness of being affected by the US policy change.

4.3.3 Heterogeneous Fixed Costs Structures across Products

The theoretical model has not differentiated across product types, but the results of this subsection suggest that the US policy spillover might exert effects on only a subset of goods. To identify the products affected by the removal of US tariff uncertainty, the analysis proceeds with estimations of subsamples which comprise only a fraction of the goods available in the baseline dataset. The first step applies equation (14) to randomly drawn product sets, comprised of 5 percent of the 3,271 products. This process is done 10,000 times, and the estiamted coefficient $\hat{b}_1$, its p-value, and the standard deviation of GAP_J in the respective sample, sd_g, are saved. Stacking together the obtained estimates results in a dataset with 10,000 estimates of b_1, which are normally distributed around the median of $\hat{b}_1^{med} = 0.704$. The constructed dataset contains also information on which good belonged to the sample, g, that produced the result.

The next step constructs a binary variable sig_{Jg}, which takes a value equal 1, when $\hat{b}_{1,g}$ is statistically significant at the 5 percent level, and zero otherwise. A probabilistic regression of sig_{Jg} on product-specific indicator variables, D_J, reveals which product is most likely to produce statistically significant results.

$$\Pr(sig_{Jg} = 1 | \mathbf{D}, sd_g) = \Phi(\beta_1' \mathbf{D} + \beta_2 sd_g + \varepsilon_{Jg})$$

The probit estimation also includes the information on the standard deviation of GAP in the respective sample and allows for cross-sample heteroscedasticity by clustering groups.[23] β_1' is

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[23]The use of product indicators implies the estimation of relative probabilities. The base-product, relative

the vector of coefficients which will inform about the probability of J being statistically more likely to contribute to significant estimates of $\hat{b}_1$. A majority of the goods (78.06 percent) produce a $\hat{\beta}_1^J$ which is positive and statistically significant at the 5 percent level.[24] These goods are coded as being elastic to the US policy change, i.e., having a fixed costs structure which favors the transmission of bilateral policy changes to its multilateral performance. Because this group is very large, a second indicator variable classifies the top 50 percent of goods as being elastic.

Table 6: China's Export to the EU after Removal of US Tariff Uncertainty, Linear Panel Estimation of Subsamples; 1995-2005

	Estimation-based		Median-based	
	Subsample (1)	Full Sample (2)	Subsample (3)	Full Sample (4)
US Tariff Threat	2.466**	0.039	3.288**	0.196**
	(0.079)	(0.073)	(0.090)	(0.071)
EU Tariff	-1.390**	-1.384**	-0.732^a	-1.355**
	(0.310)	(0.268)	(0.392)	(0.268)
EU Quota '02	0.037	0.119**	-0.0643	0.136**
	(0.040)	(0.038)	(0.054)	(0.038)
EU Quota '05	0.206**	0.183**	0.202**	0.207**
	(0.051)	(0.044)	(0.069)	(0.044)
US Tariff Threat $\times$ Elastic		0.822**		0.692**
		(0.050)		(0.043)
R-squared	0.248	0.239	0.251	0.239
Observations	217,121	284,419	139,454	284,419

Note: Table shows estimates based on Eq. (14), using alternative subsamples and specifications. Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year, included in all regressions. Robust standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

To evaluate this sorting of products, the baseline estimation is applied to the subsample of goods classified as being elastic, based on the probit estimation and on the median sample split, respectively. An alternative specification considers the full sample of products, but in-

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to which the probabilities of having $sig_{Jg} = 1$ is estimated for all other products, is therefore chosen as the specific good $\hat{J}$ which reports $sig_{Jg} = 1$ in fifty percent of its observations.

[24]Products which were in statistically significant groups in at least 57 percent of its observations always belong to this group, whereas some products qualify already with 53.7 precent of its observations with $sig_g = 1$.

teracts GAP_J^{US} with the classification of goods in the respective elasticity group. The results, shown in columns (1) and (3) of Table 6, indicate that the estimate of b_1 is now larger when only elastic products are considered. The median-based classification expectedly produces a larger point estimate. The two remaining columns (2) and (4) suggest that the interaction of GAP_J^{US} with the estimation-based group of elastic products leads to a statistically insignificant coefficient $\hat{b}_1$, whereas the median-based grouping identifies a spillover effect for the inelastic products, which is, however, smaller than the estimates obtained from any previous analysis.

Investigating further the grouping of goods into elastic and non-elastic to bilateral policy changes, leads to the question of the determinants of this elasticity. Referring to the existing literature on trade patterns of Chinese domestic enterprises and the role processing trade (Manova and Zhang, 2009; Amiti and Freund, 2010), and to the notion that goods traded within regional production networks potentially reveal different patterns, suggests that less differentiated products with a broad scope of utilization, and relatively low intensity in production should be more prone to respond to the bilateral policy change.

Table 7 presents results of probit estimations, where the dependet variable takes the value equal to one when the good is considered to be elastic with respect to the US policy change. The upper panel uses the estimation-based classification, whereas the lower panel relies on the median-based grouping, from which goods estimated to be elasitc were removed. Table 8 shows estimates on the size of $\hat{\beta}_J$, products classified as being elastic (i.e., $J = elastic$ holds). Columns (1) and (2) are based on the estimation-based grouping, while columns (3) and (4) rely on the median-based selection and excluded the estimation-based elastic products afterwards. In columns (1) and (3) the dependent variable is binary, taking a value equal one if the $\hat{\beta}_J$ is a positive outlier, and zero if it is small (and ranges in the bottom quartile). In columns (2) and (4) results are presented for a linear regression of the log of $\hat{\beta}_J$.

Table 7 suggests that differentiated products, according to the conservative classification of Rauch (1999), are less likely to appear in the samples where a statistically significant

Table 7: Multilateral Export Elasticity and Product Characteristics, Cross-sectional Estimates

Indep. Var. X	Rauch (1999)	Nunn (2007)	Input Analysis	
	Differentiated	Low Contr. Intens.	Scope	Avg. Share
	(1)	(2)	(3)	(4)
Estimation-based				
$\Pr(J = elastic\|X)$	-0.133*	0.038^a	-0.035	0.287
	(0.055)	(0.021)	(0.108)	(0.631)
Observations	3,235	3,378	1,830	1,830
Median-based				
$\Pr(J = elastic\|X)$	-0.190**	0.060**	0.077	0.171
	(0.060)	(0.023)	(0.123)	(0.728)
Observations	2,337	2,428	1,306	1,306

Note: Table shows results of probit estimation with dependent variable equal one, if product classified as elastic to bilateral policy change. Explanatory variables denoted in respective column headings. Statistical significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

effect of the US policy change on China's EU exports could be found. Homogeneous goods are accordingly more likely to be affected by the removal of tariff uncertainty. The second column indicates that the inverse of the broad measure of contracting intensity (i.e., the share of homogeneous inputs used for the production of good J) is positively related to the probability of a good being elastic to the bilateral policy change. The remaining columns of Table 7 suggest that the scope and intensity of the products' use is not related to the probability of being elastic to the bilateral policy change.[25] However, Table 8 shows that the estimated size of β_1^J, among the elastic products, is positively related with the scope of a product's utilization, weighted by the average share among production inputs. This suggests widely, but less intensively used products have a higher multilateral export elasticity with respect to bilateral tariff uncertainty.

Together, the results of this analysis are interpreted as suggestive evidence of product heterogeneity in the importance of global fixed costs, which leads to variation in their mul-

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[25]The measure of scope is computed using the Make and Use tables for the US economy in 2002, and states the fraction of industries in which a good is used as a production input. Based on the same data, the intensity of a product is measured as its average share among all production inputs in the industries where it is used for production. Details on the construction of these measures are provided on Appendix B.

Table 8: Extent of Multilateral Export Elasticity and Product Characteristics, Cross-sectional Estimates

Sample	Estimation-based		Median-based	
Model	Log-linear	Probit	Log-linear	Probit
	(1)	(2)	(3)	(4)
Differentiated	-0.247[a]	-0.028	0.143	0.020
	(0.148)	(0.018)	(0.182)	(0.019)
Low Contr. Int.	0.058	0.005	0.096	0.010
	(0.061)	(0.008)	(0.072)	(0.008)
Scope/Avg.Share	0.158**	0.022**	0.147**	0.016**
	(0.043)	(0.005)	(0.056)	(0.005)
Observations	627	1,344	392	853
R-squared	0.021	0.016	0.022	0.017

Note: Table shows results of linear and probit estimations for multilateral export elasticity on product characteristics. Statistical significance: [a] $p < 0.1$, [*] $p < 0.05$, [**] $p < 0.01$.

tilateral export elasticity. The goods most likely to be affected by the removal of US tariff uncertainty require homogeneous inputs and are homogenous themselves, while their elasticity is larger when they can be utilized for many purposes at low intensity. Given these characteristics, hypothesis of segmented exports from China across countries seems to be supported. The patterns conform to observations of Manova and Zhang (2009), where private domestic Chinese have less differentiated input sources, while they sell more goods to more destinations. The simplicity of these goods suggest that these exporters exploit their labor-cost advantage with economies of scale.

5   Conclusion

This paper analyzed a potential source of China's export boom to the EU after its accession to the WTO in December 2001. In contrast to the US, where China benefited from the establishment of permanent normal trade relations, EU trade policies had remained largely unchanged.

The theoretical model expands on previous studies of trade policy uncerainty, while it

adopts a structure that resembles characteristics of services trade models. It emphasizes the existence of a global fixed cost component which Chinese exporters must cover before they start exporting. This component can be distributed across sources of revenues, so that a firm entering a large market (the US) will find it also easier to export to third countries. The modelling strategy is motivated by an empirical literature that has assembled a number of stylized patterns concerning the export entry behavior of firms in low- and middle-income countries. Implicitly, it assumes that Chinese firms are exporters of labor services.

The theoretical analysis revealed a number of predictions which the paper submits to empirical analyses, using product-level data at the disaggregated HS6 level. The empirical strategy makes use of a difference-in-difference approach where it interacts a product specific measure of US tariff uncertainty with the event of China's WTO entry. The empirical results suggest that, indeed, there has been a positive impact of the US policy change on China's exports to the EU. The adjustment takes place at the extensive margin and occurs during a limited period of time. The size of the estimated coefficients are in reasonable orders of magnitude.

Further analyses provide supporting results that Chinese exporters face global fixed costs, which they distribute across sources of income (export destinations) in order to exploit economies of scale. However, it was also found that a meaningful redistribution of the fixed costs is initiated only, if the policy making country is large, and if its market is not part of a major regional production network. Along these lines, the paper showed that the goods most prone to be affected by the US policy change are relatively homogeneous and utilized widely.

The findings of this paper reveal an important mechanism of trade creation, which expands on the evidence documented in recent studies on the effects of trade policy uncertainty. They also suggest that (at least a fraction of) Chinese exporters actually sell labor services. In this respect, the policy design of large economies may have wide-reaching effects on both the evolution of export-oriented industries in developing countries, as well as on the intensity of

import competition in high-income countries and on export competition among low-income countries.

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A   Appendix: Robustness and Further Results

A.1   Robustness Checks

This appendix presents the results discussed in section 4.2. Table A1 shows estimated coefficients of the baseline variables and the alternative control variables, described in the main text of the paper. Details on the construction of each control variable are provided in Appendix B. Table A2 further presents an overview of how the resulting measures are correlated with each other.

Table A1: Chinese Exports to the EU-15 and the Removal of US Tariff Uncertainty; Robustness analysis, 1995-2005

	Unobserved productivity shocks				Effects of Chinese Policies					European Events		Full Set
	Δ China's RCA	× X-rate	Qual.-adjusted Price		Contracting Intensity		Imported Inputs		Subsidies	RCA^{GIPS}	Japan	
			Export	Export/Import	Strict	Broad	Δ Tariffs	Δ Value				
Dependent variable ln R_{Jnt}	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
US Tariff Threat	0.503**	0.511**	0.796**	0.789**	0.609**	0.582**	0.559**	0.577**	0.680**	0.670**	0.556**	0.555**
	(0.065)	(0.065)	(0.069)	(0.069)	(0.067)	(0.066)	(0.067)	(0.067)	(0.066)	(0.066)	(0.177)	(0.069)
Applied Tariff	-1.002**	-0.996**	-1.463**	-1.477**	-1.416**	-1.539**	-1.412**	-1.419**	-1.454**	-1.548**	-1.498**	-1.023**
	(0.263)	(0.263)	(0.268)	(0.268)	(0.266)	(0.266)	(0.266)	(0.265)	(0.266)	(0.265)	(0.687)	(0.266)
EU Quota '02	0.201**	0.191**	0.083*	0.092*	0.098**	0.098**	0.105**	0.099**	0.109**	0.162**		0.220**
	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)		(0.038)
EU Quota '05	0.223**	0.225**	0.165**	0.172**	0.177**	0.187**	0.177**	0.169**	0.187**	0.148**		0.196**
	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)	(0.044)		(0.044)
ΔRCA (CN)	1.990**											1.864**
	(0.055)											(0.058)
ΔRCA (CN) (×X-rate)		0.203**										
		(0.006)										
Qual.-adj. Price			-0.484**									-0.119*
			(0.059)									(0.059)
Qual.-adj. Price (deflated)				-0.482**								
				(0.048)								
Contr. Intensity (strict)					0.322**							0.423**
					(0.061)							(0.059)
Contr. Intensity (broad)						0.758**						
						(0.094)						
Δ Tariff China							-0.522**					-1.073**
							(0.182)					(0.179)
Δ Imports China								0.001				
								(0.016)				
Subsidies 1 (ADF)									0.175**			0.087*
									(0.039)			(0.040)
Subsidies 2 (CVD)									0.226^{a}			0.340**
									(0.132)			(0.128)
RCA (GIPS)										0.115**		0.069**
										(0.007)		(0.007)
Observations	288,693	288,693	287,405	287,405	290,061	290,061	289,652	289,652	290,761	290,761	31,606	284,533
R-squared	0.245	0.245	0.237	0.237	0.234	0.235	0.235	0.235	0.234	0.235	0.269	0.250

Note: Table shows estimates based on Eq. (14), including control variables as indicated in column headings. Fixed effects: Jn=product-destination, nt=destination-year, included in all estimations. Column (11) shows results of estimating Chinese exports to Japan, and the associated applied tariff rates. Robust standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

Table A2: Chinese Exports to the EU-15 and the Removal of US Tariff Uncertainty; Robustness analysis, 1995-2005

		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
US Tariff Threat	(1)	1.00											
ΔRCA^{CN}	(2)	0.04	1.00										
$\Delta RCA^{CN} \times$X-rate	(3)	0.02	0.40	1.00									
Qual.-adj. Price	(4)	0.28	-0.11	-0.04	1.00								
Qual.-adj. Price, deflated	(5)	0.27	-0.09	-0.04	0.91	1.00							
Contr. Intensity (strict)	(6)	0.18	-0.09	-0.04	0.33	0.33	1.00						
Contr. Intensity (broad)	(7)	0.20	-0.09	-0.04	0.14	0.13	0.66	1.00					
Δ Tariff CN	(8)	-0.10	0.03	0.01	-0.05	-0.02	-0.09	-0.26	1.00				
Δ Imports CN	(9)	-0.00	-0.09	-0.03	-0.09	-0.13	-0.16	0.09	-0.71	1.00			
Subsidies (ADF)	(10)	-0.04	0.06	0.02	-0.02	-0.02	-0.04	0.01	0.05	-0.03	1.00		
Subsidies (CVD)	(12)	-0.02	0.05	0.02	-0.02	-0.01	-0.02	-0.01	0.03	-0.01	0.32	1.00	
RCA^{GIPS}	(13)	0.08	0.22	0.09	0.01	0.03	-0.04	-0.04	0.02	-0.13	0.04	-0.01	1.00

Note: Table shows coefficients of correlations for variables used in for estimation of (14), in alternative versions.

Figure A1 shows the estimated coefficients and confidence intervals of the US policy spillover, $\hat{b}_1$, across samples where the respective European destination, indicated on the vertical axis, has been removed.

Figure A1: Chinese Exports to the EU-15 and the Removal of US Tariff Uncertainty; Destination-specific variation, 1995-2005
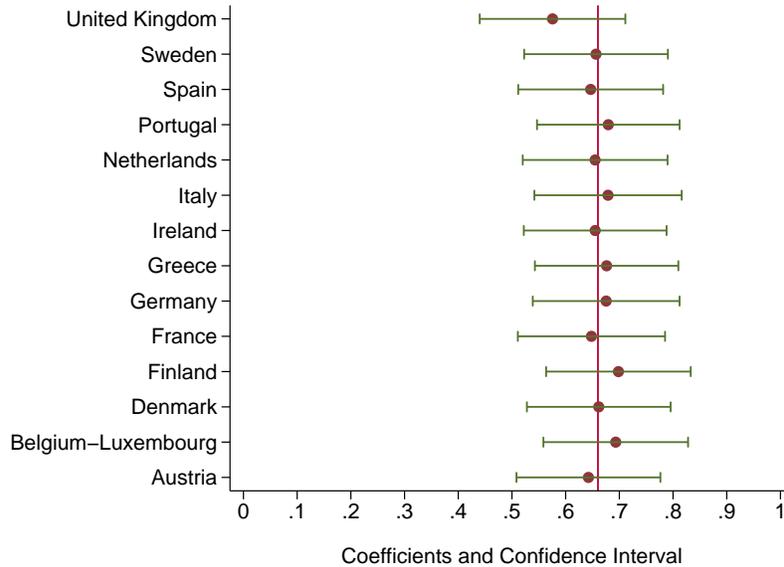
A.2 Further Results

Table A3 presents results of the baseline specification for an increased sample period, up to the year 2012. Columns (2) and (3) adopt a dynamic specification, which includes the

lagged dependent variable as an additional explanatory variable. The remaining columns (4) through (6) present results for the effect on the normalized growth rate of exports for alternative sample lengths. The discussion of these results is presented in Section 4.3.1 of the paper.

Table A3: Chinese Exports to the EU-15 after the Removal of US Tariff Uncertainty; Static and Dynamic Effects, 1995-2012

	Levels 1995–2012			Growth 1995–T		
	Baseline	FE	POLS	$T = 2002$	$T = 2005$	$T = 2010$
	(1)	(2)	(3)	(4)	(5)	(6)
US Tariff Threat	0.748**	0.216**	0.332**	0.441**	0.084**	-0.011
	(0.075)	(0.048)	(0.0186)	(0.067)	(0.030)	(0.021)
EU Tariff	-0.444^a	-0.547**	0.542**	0.286	0.466^a	-0.234
	(0.253)	(0.203)	(0.107)	(0.326)	(0.249)	(0.173)
EU Quota '02	-0.082*	0.005	-0.030**	0.296**	0.138**	0.056**
	(0.041)	(0.026)	(0.0107)	(0.039)	(0.018)	(0.013)
EU Quota '05	0.346**	0.282**	0.0259*		0.403**	0.122**
	(0.043)	(0.027)	(0.0121)		(0.035)	(0.011)
EU Quota '09	0.977**	0.570**	0.266**			0.011
	(0.059)	(0.034)	(0.0213)			(0.015)
Lagged Exports		0.432**	0.844**			
		(0.003)	(0.00114)			
Observations	561,498	474,266	474,266	192,739	298,789	588,356
R-squared	0.345	0.488	0.786	0.023	0.017	0.0323
Fixed effects	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St

Note: Table shows estimates from alternative specifications. Columns (1) through (3) compare the baseline specification to dynamic panel estimates providing lower and upper bounds of the coefficient for lagged exports (shown in the last row). Columns (4) through (6) analyze the transitional growth effect of $GAP_{J,99}^{U.S.}$ by comparing periods of different length after the policy change. Abbreviations represent fixed effects model (FE), pooled OLS (POLS), and the last year observed in the sample for which results are shown (T). Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year. Standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

Table A4 presents additional results for inferring the existence of fixed costs redistribution. The main variable of interest is Δs_{Jn}^{CN}, the change in the share of China's product J exports shipped to destination n, with ns denoted in the respective column headings. The measure serves as a proxy for $\Delta\theta_{Jn}$. The discussion of these results is presented in the main text of the paper, in Section 4.3.

Table A4: Chinese Exports to the EU-15 and relocation of global fixed cost shares; 1995-2012

	Canada	Japan	New Zealand	South Korea
	(1)	(2)	(3)	(4)
Δs_{Jn}^{CN}	0.486^a	-0.373**	-0.324	-0.359**
	(0.263)	(0.052)	(0.965)	(0.096)
EU Tariff	-1.388**	-1.339**	-1.401**	-1.393**
	(0.266)	(0.266)	(0.266)	(0.266)
EU Quota '02	0.138**	0.136**	0.137**	0.137**
	(0.037)	(0.037)	(0.037)	(0.037)
EU Quota '05	0.190**	0.179**	0.190**	0.190**
	(0.044)	(0.044)	(0.044)	(0.044)
Observations	290,216	290,216	290,216	290,216
R-squared	0.233	0.234	0.233	0.233
Fixed effects	All	All	All	All

Note: Table shows estimates of China's exports to EU-15, conditional on change in fraction of product J exports, Δs_{Jn}^{CN}, to destinations denoted in column headings. Fixed effects: Jn=product-destination, nt=destination-year, and st=sector-year, included in all specifications. Standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

B Dataset Construction

This appendix describes the datasources and procedures used to construct the dataset for the paper *US Policy Spillover(?): China's Accession to the WTO and Rising Exports to the EU*. Table B1 presents the primary sources, their resulting variables in the dataset, and the names of the STATA *do-files*, in order of their appearance. The following paragraphs provide details on their content.

Table B1: Data Files, Programs, and Sources

No.	name of do-file	primary sources	variables
1	1_data_us_tariffs_gap.do	Feenstra et al. (2002): http://www.nber.org/data/	gap1 gap2 gap3
2	2_data_eu_tariffs.do	World Integrated Trade Solutions (WITS): http://wits.worldbank.org/ Council-Regulation (EC) No 3281/94; 2820/98; 2501/2001; 2211/2003; 980/2005; 732/2008	EUtar1 EUtar2
3	3_data_eu15_imports.do	UN Comtrade Online Database: un.comtrade.org	cn2eu
4	4_data_eu15_mfaquotas.do	Système Intégré de Gestion de Licenses: http://trade.ec.europa.eu/sigl/	EUquo02 EUquo05 EUquo09
5	5_data_contracting_intensity_nunn2007.do	Nunn (2007): http://scholar.harvard.edu/nunn/pages/data-0	nunn_strict nunn_broad
6	6_data_quality_adjusted_prices.do	Feenstra and Romalis (2014): http://cid.econ.ucdavis.edu/Html/Quality_Data_Page.html	QAXP1 QAXP2
7	7_data_rca_south.do	UN Comtrade Online Database (see no. 3)	RCA_GIPS
8	8_data_exports_to_japan.do	WITS (see no. 2)	JPtar cn2jp
9	9_data_export_subsidies.do	Bown (2014, 2015): Global Countervailing Duties Database (GCVD); Global Antidumping Database (GAD)	ADF CVD
10	10_data_china_rca.do	Centre d'Études Prospectives et d'Informations Internationales (CEPII) http://www.cepii.fr/CEPII/en/bdd_modele/presentation.asp?id=26	RCA0 RCA1
11	11_data_china_xRate.do	Penn World Tables 8.0: http://www.rug.nl/research/ggdc/data/pwt/	EUxr USxr
12	12_data_share_in_CN_exports.do	UN Comtrade Online Database (see no. 3)	USs NZs
13	13_data_CN_to_US_Exports_and_Tariffs.do	WITS (see no. 2); Schott (2008): http://faculty.som.yale.edu/peterschott/sub_international.htm	UStar cn2us
14	14_data_CN_inputs_tariffs.do	WITS (see no. 2); BEA Make&Use Tables http://www.bea.gov/industry/io_benchmark.htm	cnINtar cnINimp
15	15_data_dataset_final.do	files no. 1-14	
16	16_data_product_usage_scope_and_bec.do	BEA Make&Use Tables (see no. 14)	mshare scope
17	17_data_rauch1999_hs6.do	Rauch (1999): http://www.macalester.edu/research/economics/PAGE/HAVEMAN/Trade.Resources/TradeData.html	rauch_con rauch_lib
18	18_data_charges_trend.do	Schott (2008) (see no. 13)	charge

1 US MFN and Column 2 Tariffs. The raw tariff dataset is downloaded from the source stated in table B1, and is used to construct 3 alternative measures of the US tariff threat, GAP_J^{US}. Each of the measures represents numbers as observed for the year 1999. The first measure, the variable 'gap1', is the one used in all regressions for which results are presented in main text of the paper. The two alternative measures are constructed as suggested by Pierce and Schott (2013) and Handley and Limão (2013), i.e., 'gap2' and 'gap3', respectively. 'gap2' is calculated as the simple difference between the Column 2 and the MFN rate, i.e., gap2$= \tau_{hts}^{col2} - \tau_{hts}^{mfn}$. The equation for 'gap3' is gap3$= 1 - (\tau_{hts}^{col2}/\tau_{hts}^{mfn})^{-\sigma}$, where $\sigma = 3$.

Before any of these measures were calculated, missing information on MFN and Column 2 ad-valorem equivalent rates (mfn_ave and col2_ave) was filled with the respective values on ad-valorem tariffs. Products with missing information in both cases were discarded. In cases where Column 2 rates were smaller than the MFN rate, specific Column 2 rates were used (col2_spec), if they were higher than the MFN rate. The remaining Column 2 rates below MFN rates were set equal to the MFN rate to ensure that the tariff threat cannot be negative. After computing 'gap1'-'gap3', the measures were aggregated to the 6-digit HS6 level and converted to the HS1988/1992 classification. The resulting dataset used for the paper is named `1_data_us_gap_1999.dta`. A second dataset, produced by the do-file, contains annual values of 'gap1'-'gap3', for the period 1989-2001 (`1_data_us_tariffs_gap.dta`).

2 EU Tariffs. The two tariff variables, 'EUtar1' and 'EUtar2', are compiled in three steps. The former is the one used in the analyses for which results are presented in the paper.

Step 1 identifies products which are subject to potential (or actual) exceptions from the European GSP, as published in the respective Council Regulations. To trace these goods over time, correspondence tables from Bernard et al. (2012) and from the EUROSTAT RAMON database are used. At the end of step 1, each good carries information on whether it is subject to GSP discounts on MFN rates, how much this discount is, and to what extent China benefits from this discount.

The second step uses information on EU tariff schemes, accessible through the WITS

databases. The schedules report applied MFN, GSP, and several special rates for each year and product at the CN-8 level. The product codes in this database are not throughout consistent with the announced re-classifications, so that old codes may continue to appear in later years. Accordingly, also here, GSP and MFN rates are mapped to their actual CN codes applicable in the respective year. At the end of step 2, the tariff rates stated in the European tariff schedules are matched with the information on GSP discounts, obtained from step 1.

Step 3 constructs variables from this information: Applied MFN and GSP rates, as reported in the tariff schedules (ad-valorem equivalents), and applicable GSP rates and GSP tariffs for China, according to the Council Regulations. The four variables are aggregated to the 6-digit HS6 level and matched with direct information on EU imports and applied tariffs on Chinese imports (also obtained from WITS). With the fifth measure at hand, all HS6 codes are converted to the HS1988/1992 classification, so that applicable and applied tariffs on Chinese products can be observed for each product-year, independent from whether trade had actually occurred or not.

Finally, two measures are kept in the dataset: the first measure, 'EUtar1', denotes the applied tariff on Chinese imports, as stated by the primary source with observed trade, and the applicable rate when trade did not occur. The second measure, 'EUtar2', denotes the potentially applicable tariff rate under the assumption that China has graduated from GSP rates for all products envisaged by the respective Council Regulations. The log difference between EUtar2 and EUtar1 would constitute a measure of tariff uncertainty in the EU. The final output file is named `2_data_gsp_hs6.dta`.

3 EU-15 Imports. EU-15 imports measure the value of goods European countries imported from China, including transportation costs (i.e., CIF values). The raw data is disaggregated at the 6-digit HS level. The first step aggregates the partial files for each EU country into a destination-specific file, containing information on imports from all countries across years. The second step harmonizes country names for reporting and partner countries. In the third step, product codes are harmonized over time, so as to reflect HS6 codes of the

HS1988/1992 classification, and computes absolute import values of Chinese goods, as well as China's import market share. The final step sticks together the destination-specific information on Chinese imports, and aggregates the values imported by Belgium and Luxembourg (the countries report trade statistics separately only since 1999). A fully balanced panel is constructed to denote missing values as inactive trade relationships. The resulting variable, 'cn2eu', denotes the dependent variable in all regressions for which results are presented in the paper. The do-file produces a dataset named `3_data_eu15_imports.dta`.

4 EU Import Quotas on Chinese Goods. The removal of import quotas is indicated by variables which take values between 0 and 1 for the respective product-years when the restrictions were lifted. This occured three times, in 2002, 2005, and 2009. The raw dataset, containing information on quotas in the textiles, clothing, and apparel industries reports the amount of a good, each country that was constrained by MFA/ATC restrictions could export in a respective year. Moreover, percentages of quota utilization are reported.

The first step selects the information about China and harmonizes the coding of product categories, using the descriptions available from the SIGL category guide. Based on the percentage of quota utilization level by EU member states, the categories to which quotas were applied are identified. A secondary measure (not used in this paper) identifies those products where the quota was binding, i.e., when the utilization rate was equal to 75% or higher, on average. For each of the two measures, indicator variables are constructed, denoting the year where the quota was lifted. In a final step, the product categories are mapped to the corresponding codes of the Combined Nomenclature. These codes are aggregated to the HS6 level and concorded to the HS1988/1992 classification. For each quota variable, HS6 products may encompass categories with both expiring and non-expiring quotas. Accordingly, the variables 'EUquo02'-'EUquo09' are to some extent continuous between 0 and 1. The resulting dataset is named `4_data_EU_MFA_hs6.dta`.

5 Contracting Intensity. This variable is used in the robustness checks section, and serves to investigate whether the liberalization of Chinese foreign investment regulations have influenced the results for the US policy spillover. To utilize this measure, it is neccesary to identify the exact categories of the NAICS classification. For instance, a six-digit industry 1111A0 (Oilseed Farming) maps to the codes 111110 and 111120, whereas 1111B0 contains (Grain Farming) contains the remaining categories 111130 through 111190. After this decomposition, detailed NAICS codes are matched with the measures of contracting intensity from Nunn (2007). The second step uses the concordance tables from Pierce and Schott (2009), in order to map the measures to the HS6 products used in the final dataset. 291 NAICS industries can be assigned to 4,120 HS6 products. The two obtained measures of contracting intensity differ in that 'nunn_strict' < 'nunn_broad'. The output file is named `5_data_contracting_intensity_nunn2007.dta`.

6 Quality-adjusted Export Prices. Quality-adjusted export prices are reported for 4-digit SITC product categories, differentiating alternative measurement units of a category, and in relative terms in each product-year. The base unit changes across products and years. Therefore, for each SITC product-unit-year, the reported quality-adjusted price of a country ('qa_price_exp') is deflated by the respective group average, expressing relative quality-adjusted prices. The relative prices eliminate measurement units so that, for each SITC category, quality-adjusted prices are aggregated as weighted averages of the unit-specific values.

Instead of using annual information on these prices, the second step calculates the ratio of average quality-adjusted export prices in 2002-2004 relative to 1997-1999, i.e., before and after China's WTO entry. Analogously, quality-adjusted import prices are calculated for the European destination countries. The final step maps SITC codes to the appropriate HS6 product categories. The resulting measures, 'QAXP1' and 'QAXP2', reflect China's quality-adjusted export price, and the same price, deflated by the European import price. The resulting dataset is named `6_data_QadjPrices.dta`.

7 Revealed Comparative Advantage of South-European Economies. The RCA measure for South-European countries encompasses Greece, Italy, Portugal, and Spain (GIPS). Instead of using a direct measure for global exports, this 'RCA_GIPS' variable measures the revealed comparative advantage in EU markets. Hence, the data on imports from the UN Comtrade database is processed as described in paragraph 3 (EU Imports) to obtain a balanced panel with information on each GIPS country's exports to the rest of the EU-15. Based on the information on imports from all other countries, the RCA is calculated as

$$RCA_{Jt}^{GIPS} = \frac{x_{Jt}^{GIPS}/X_t^{GIPS}}{x_{Jt}/X_t},$$

where $X \equiv \sum_{J=1} x_J$, and the ratio in the denominator indicates rest of EU's imports from rest of the world. The measure is finally averaged across values for the three years before China's WTO entry, 1999-2001. The resulting dataset is named `7_data_rca_south.dta`.

8 Japanese Imports. Data on Japanese imports from China and applied tariffs are accessed through the WITS. The data readily reports imports at the 6-digits HS level and applied tariff rates. The tariffs applicable for the cases where no trade was observed are obtained from Japan's tariff schedules, reporting MFN and GSP rates. Whenever GSP rates were not reported for a good, but so were MFN rates, the latter was used to fill missing information on applicable tariffs. The resulting variables 'cn2jp' and 'JPtar' are stored in a file named `8_data_exports2jpn.dta`.

9 Chinese Export Subsidies. Export subsidies are inferred using information filings regarding countervailing duties and anti-dumping investigations. The former, constituting the variable 'CVD' can be observed only for 35 products, in the years 2010 or later. For these products, 'CVD' is set equal to 1.

Anti-dumping (AD) investigations, in turn are selected for cases where the final decision on domestic industry injury was affirmative ('f_inj_dec'="A"). Of these cases, observations

with missing dates on revocation of AD measures ('revoke_date'="MI", ".") are discarded. Products are matched with HS6 codes, and dates of approval and revocation are aggregated to annual observations. Finally, AD cases initiated after 2003 are selected, so that 140 HS6 products have potentially benefitted from export subsidies, around the date of China's WTO entry. The resulting dataset, including both countervailing duties, 'CVD', and anti-dumping filings, 'ADF', is named `9_data_export_subsidies.dta`.

10 China's Revealed Comparative Advantage. The variables 'RCA0' and 'RCA1' denote China's revealed comparative advantage, as reported in the CEPII RCA data. 'RCA0' denotes the average value for the years 1995-1997, whereas 'RCA1' refers to the period 2008-2010. The dataset, named `10_data_rca_china.dta`, provides this information for 968 product groups, aggregated at the 4-digit HS level.

11 Chinese Exchange Rate. This program selects China's market exchange rates in local currency units per US dollar ('xr'), which was pegged between 1998 and 2004. This rate is divided by the respective exchange rates of the EU-15 member states, to obtain cross-rates expressing China's Yuan in units of the European currencies. Both rates 'xr_us' and 'xr_eu' are saved in the file `11_data_xrate_china.dta`.

12 Destination Share in Chinese Exports. This program uses information on exports, reported by China at the 6-digit HS level, to compute the fractions of China's exports of product J shipped to destination n in year t, s_{Jnt}^{CN}. Averages of this fraction are computed for the periods 1992-1996, $T = 0$, and 2004-2008, $(T = 1)$. The change of this share, $\Delta s_{JnT}^{CN} = s_{JnT}^{CN} - s_{JnT-1}^{CN}$, indicates the relative increase of China's product J exports to country n. The values for $n = \{CA, EU, JP, NZ, KR, US\}$ are selected to compare the allocations of exports countries of different size. The resulting dataset is named `12_data_share_in_CN_exports.dta`.

13 US Imports and Tariffs. Applied MFN tariffs by the US, reported at the 6-digit HS level, are concorded to the HS1988/1992 classification. This information is matched with data on US imports from China. The resulting dataset contains a balanced panel of US imports and tariffs for Chinese goods, during the period 1998-2006, and is named `13_data_CN_to_US_Exports_and_Tariffs.dta`.

14 Tariffs and Imports of Production Inputs. Chinese tariffs on and imports of HS6 products J are concorded to the HS1988/1992 classification. Tariffs are expressed as $\tau = 1 + (\%rate/100)$. The annual observations are averaged over China's the pre- and post-WTO entry periods, 1996-2000 vs. 2001-2005, and log differences of each variable are computed to obtain the relative increase of applied tariffs and imports from the rest of the World.

The second step constructs a correspondence file, which is then used to map China's tariffs and imports to the respective NAICS industries, as defined by Pierce and Schott (2009) for the year 1995. At the end of this step, the amount and tariffs on Chinese production inputs are observable at the 4-digit NAICS industry level.

Using the "Make and Use Tables" of the US economy for the year 2002, Chinese inputs are matched with the associated output industries. The information on NAICS input tariffs and imports is aggregated to the associated output industries. For example,

$$\Delta\tau_{OUT} = \sum_{IN} w_{IN,OUT}\Delta\tau_{IN},$$

where w denotes the percentage of input IN in total inputs required by the output industry OUT. The resulting variables inform about the relative change of China's input tariffs and imports in the respective output industry.

In the final step, the NAICS output industries are mapped back to HS6 products. The resulting dataset provides information on tariff and import changes for production inputs, based on about 50 manufacturing industries, at the HS6 product level J. The dataset is named `14_data_imported_inputs`.

15 Final Dataset. This do-file merges all the information assembled during the previous steps, 1-14, and assigns the ultimate variable names, as reported in table B1. In addition, sectorial groups, indicating sectors, S, are defined. The output file is named `15_dataset_final.dta`. The following 2 steps, 16 and 17, describe the construction of additional data which was used to obtain the results at the end of section 4.3 of the paper.

16 Product Usage and Scope. This program constructs indicators for the scope and intensity of Chinese exports as inputs in other sectors. As in the previous paragraph, HS6 products are matched with the respective NAICS input industries. For each input industry, IN, its scope of usage is computed:

$$scope_{IN} = \frac{1}{N} \sum_{OUT} D_{IN,OUT},$$

where $D_{IN,OUT}$ equals one, if IN is used in production of OUT, and zero otherwise. N denotes the total number of output industries.

Analogously, the average intensity of each input industry IN is computed as

$$mshare_{IN} = \frac{1}{N_{IN}} \sum_{OUT} s_{IN,OUT},$$

where s denotes IN's fraction of total input values, and N_{IN} measures the number of industries where IN is used.

The resulting measures, 'mshare' and 'scope' , are finally mapped back to their associated HS6 product categories J.

In addition to these measures of product usage, HS products J are matched with associated Broad Economic Categories (BEC), which differentiate between industrial supplies, fuels and lubricants, capital goods, transport equipment, consumer goods, and various subcategories. The resulting dataset is saved as `16_data_product_usage_scope_and_bec.dta`.

17 Differentiated and Homogeneous Goods. The Rauch (1999) is reported at the 4-digit SITC Rev.2 level, so that it has to be matched with HS6 codes from correspondence table of the United Nations Statistics Division (UNSD). Both the conservative and the liberal classification are taken into consideration. In the resulting dataset, products are by an indicator variable equal 1, if it is classified as being differentiated, e.g. 'con'="n". The indicator variable equals zero in all other cases, so that reference-priced goods and those traded on organized exchanges are considered to be homogeneous goods. The information is saved in the dataset named `17_data_rauch1999_hs6.dta`.