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

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Abstract

The paper explores the causes of China's rising manufacturing exports to the EU after WTO accession. While the European trade policy environment remained largely unchanged in most sectors, a spillover from a change in US trade policies towards China is emphasized. In the proposed model the transmission occurs through a global component of the fixed costs firms must pay in order to export. If a large fraction of this component can be covered from exporting to one destination, exporters will serve also other markets to maximize their profits. The empirical analysis makes use of the removal of US tariff uncertainty in conjunction with China's WTO accession. It shows 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. The results have implications for the scope of international policy negotiations and provide suggestive evidence on the nature of the fixed costs that manufacturing firms in low-wage countries must overcome.

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

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1 Introduction

When China entered the World Trade Organization (WTO) in December 2001, industrialized countries saw an increasing amount of Chinese goods flowing into their domestic markets. Since then, numerous studies have been investigating the consequences of the increased competition on domestic firms and workers. However, little is known about the fundamental causes of China's export boom. Import duties levied on Chinese products mostly corresponded to preferential rates even before WTO entry. This rules out tariff reductions as the most obvious explanation in standard trade models.[1] In search of alternative causes, recent studies appealed to the idea of trade policy uncertainty (TPU). Several papers show that a US policy change explains well the timing and structure of China's export boom to the US (Handley and Limão, 2013; Feng et al., 2014). It also conforms to sectoral patterns observed in the decline of US manufacturing employment since 2001 (Pierce and Schott, 2013).

While the US policy change towards China reconciles the evolution of their bilateral trade relationship, it does not explain the surge of Chinese exports to other high-income countries. In particular, the EU experienced a similar increase of Chinese goods flowing into its markets. This is surprising, because EU trade policies towards China were less uncertain and did not change in the way they did in the US. China's tariff status in the EU was governed under the Generalized Scheme of Preferences (GSP). Since the 1980s, it grants preferential tariffs below most-favored-nations (MFN) rates, and incorporates a transparent, performance-based graduation mechanism. Inspecting the applicable EU tariffs on Chinese products since the mid-1990s suggests that they decreased slower than those for most other trade partners. Shortly after China's WTO entry they actually increased due to graduation from the preferential GSP rates.[2] Nevertheless, Figure 1 shows that, after WTO entry,

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[1]Frequently used procedures were, thus, using China's export structure observed in other destinations as an instrument for domestic market penetration (e.g. Autor et al., 2013; Dauth et al., 2014). Some studies exploit the removal of quotas in the textile and clothing industry (e.g. Brambilla et al., 2010; Utar, 2014).

[2]This pattern is revealed by EU tariff schedules available at the World Integrated Trade Solutions (WITS) databases and from the several European Council regulations of the GSP in this period.

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China's exports to the EU departed from its long- and medium-run trend.[3]

Figure 1: Real Commodity Exports from China to EU-15 Countries; Log-scale; 1962-2011
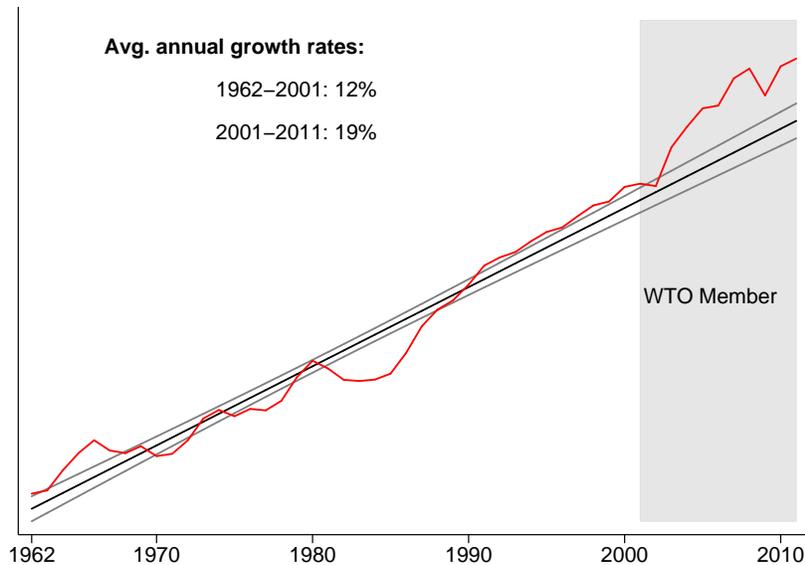
Note: Author's calculations based on NBER-UN World Trade Flows, UN Comtrade, and Penn World Tables 8.0. Exports expressed in international dollars (PPP 2005). Fitted line shows trend for 1962-2001 and 95% confidence intervals.

This paper explores the possibility that the US policy change towards China encouraged Chinese exports to the EU. In doing so it extends existing work where the mechanism through which bilateral TPU operates provides no explanation of a spillover to multilateral trade. Feng et al. (2014) suggest that the removal of US tariff uncertainty faced by Chinese exporters facilitated market entry through a reduction of the expected tariff rate. This is also suggested by Handley and Limão (2013) who, as an additional mechanism, consider technology upgrades by firms that exported to the US before the policy change. However, since such technology upgrades are modelled to reduce distribution costs in the destination, the effect is limited to their bilateral setting. To generate a spillover, this paper considers an alternative channel. Similar to the exposition of Hanson and Xiang (2011), it assumes that Chinese firms willing to export must incur both a country-specific and a global fixed cost. Empirical evidence

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[3]Throughout this paper, the European Union will be referred to as the EU-15; i.e. the countries constituting the EU between 1995 and 2004.

from Iacovone and Javorcik (2012) justifies the assumption of a global fixed cost for low-wage countries. They observe that Mexican firms upgrade their products prior to becoming exporters.

The proposed model suggests that a change in US policies towards China lowers the productivity threshold at which its firms can profitably export to third countries, given that they serve the US, too. Generally, the effect is larger when the policy-making economy is able to cover a major fraction of the global fixed cost. Bilateral fixed costs ensure that more productive firms export to a larger number of destinations, a stylized pattern in manufacturing trade data (e.g. Eaton et al., 2004). A reduction of TPU in the US implies that China increases its exports to the EU. This occurs at the extensive margin through entry of firms into new destination markets. To test the predictions empirically, the paper uses disaggregated data on Chinese exports in the period 1995-2005. The impact of the US policy change is evaluated via a difference-in-difference (DID) strategy. It exploits cross-product differences of the US "tariff threat" under TPU, and compares the periods before and after China's WTO entry. The results reveal a robust and positive impact of the US policy change on Chinese exports to the EU. The estimated coefficients appear to be in reasonable orders of magnitude, considering model parameter values used in the literature. It is also confirmed that trade increases at the extensive margin, i.e. through the creation of new trade relationships at the product-destination level. Further analyses provide suggestive evidence in support of a redistribution of global fixed costs. An extended sample period up to the year 2012 reveals that the effect of the US policy change phases out after a few years.

The paper makes several contributions to the literature. It shows 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 observed in bilateral studies (Handley, 2012; Handley and Limão, 2012, 2013). The proposed transmission channel, a global fixed cost component, generalizes the findings of Hanson and Xiang (2011), who focussed in their analysis on services exports. Although global fixed costs seem to contradict

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stylized patterns where firms enter markets one by one, a reason for why this could apply in the presented case is the large amount of processing trade in China (Amiti and Freund, 2010). It suggests that firms do not sell their goods directly to consumers but rather produce for firms in high-income countries which then place the good on the market. Chinese firms pay fixed costs to meet general standards required by those firms, irrespective of their provenance. Country-specific fixed costs remain an important feature of manufacturing trade, as they influence the costs of moving goods between source and destination countries. Finally, as product-specific trade policies and uncertainty contribute to the formation of trade patterns and investment, they might affect the development of seemingly random patterns of specialization observed across similar countries (Hausmann and Rodrik, 2003).

The paper proceeds as follows. Section 2 describes the trade policy environments faced by Chinese exporters, and argues that TPU was higher in the US than in the EU. Section 3 introduces the theoretical model, which links bilateral policies to a multilateral dimension and derives the testable predictions. Section 4 explains the empirical strategy and the data used to carry out the analysis. Section 5 presents and discusses the results. Section 6 concludes.

2   US and EU Trade Policies towards China

2.1   China-US Trade Relations

In the late 1970s the US and China established their diplomatic relations. In 1980 the US granted China preliminary MFN status for its exports. Prior tariffs corresponded to the "Column 2" schedule the US typically applies to non-market economies. These were originally defined under the Smoot-Hawley Tariff Act of 1930. In many cases Column 2 rates are much higher than the MFN rates, which were gradually dismantled during GATT/WTO negotiations. The preliminary nature of China's MFN status in the US entailed the risk that it would return to apply Column 2 rates.

Approval of MFN rates for China required a majority of votes in the US Congress, and

guaranteed the status for one additional year. Accordingly, Chinese exporters could be certain about applied tariffs in the present year but not for those that would follow. Handley and Limão (2013) quote a number of business practitioners and politicians suggesting that this form of TPU deterred investments into Chinese exports to the US. Moreover, in the 1990s it was witnessed that China's MFN status was close to being overturned. In the aftermath of the Tiananmen Square incident, in 1989, political opposition to China's MFN status arose, pointing out the violation of human rights standards. Pierce and Schott (2013) emphasize that, in the early 1990s, 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 amounted to 38 percent, on average. During these years, 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 its collision with a Chinese fighter jet over the South China Sea. It was disassembled and returned to the US after several diplomatic interventions.

The entry of China into the WTO was decided in December 2001 and has been effective since January 2002. Upon this event the US granted China "permanent normal trade relations" (PNTR). This removed the inhibiting effect of TPU for Chinese exporters and encouraged their entry into the US market. In particular, the fact that the preliminary MFN status was never actually overturned makes this policy change appropriate for consideration as a natural experiment. Its evaluation requires information about the US Column 2 and MFN tariff schedules for the years prior to China's WTO accession. Figure 2 shows how these rates differ according to the US tariff data that is going to be used in this paper. It shows that MFN rates gradually decline to below 5% for an average manufacturing product, while Column 2 rates increased from 34 to 38 percent. The log difference between the two rates, calculated for 1999, suggests that the threat of tariff increases was present across all industries.[4]

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[4]In Panel (b) of Figure 2 the threat is calculated as $\ln \tau^{Col} - \ln \tau^{MFN}$; a tariff of 5% implies $\tau = 1.05$.

Figure 2: Average US MFN and Column-2 Tariffs 1989-2001, and the Tariff Threat towards China

(a) US Column-2 and MFN Tariff (%)

(b) Log-difference across Sectors, 1999
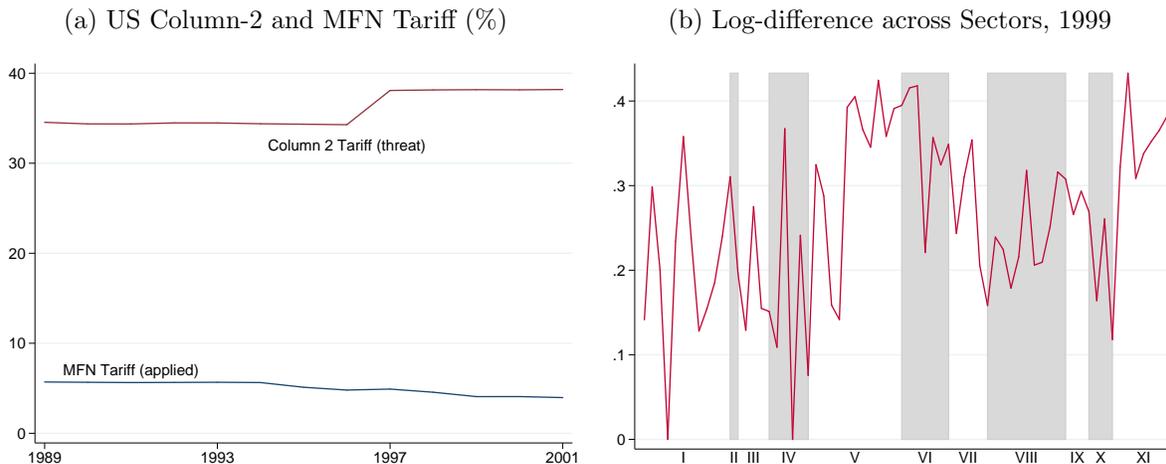
Note: Author's calculations based on US Tariff Data (Feenstra et al., 2002). Horizontal axis of Panel (b) denotes manufacturing sectors and their sub-chapters: I Chemicals, II Plastic/Rubber, III Hides/Leather, IV Wood Products, V Textiles, VI Apparel/Footwear, VII Stone/Glass, VIII Metals, IX Machinery/Electronics, X Transport, XI Other Manufactures.

2.2 China-EU Trade Relations

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 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 TPU, as it sets out tariff preferences for several years. The GSP also includes a graduation mechanism which implies that a country may return to MFN rates. This happens when it reaches a certain level of economic development, or when it becomes a dominant exporter of a good in comparison to other GSP beneficiaries. These criteria are transparent so that Chinese exporters should have faced less uncertainty regarding future tariffs. When China entered the WTO, nothing changed in these formal procedures.

Figure 3 depicts European MFN and GSP tariffs, as well as those applicable to China for

Figure 3: Average European Union Applied Tariffs (%); Ad-valorem Equivalents, 1995-2012
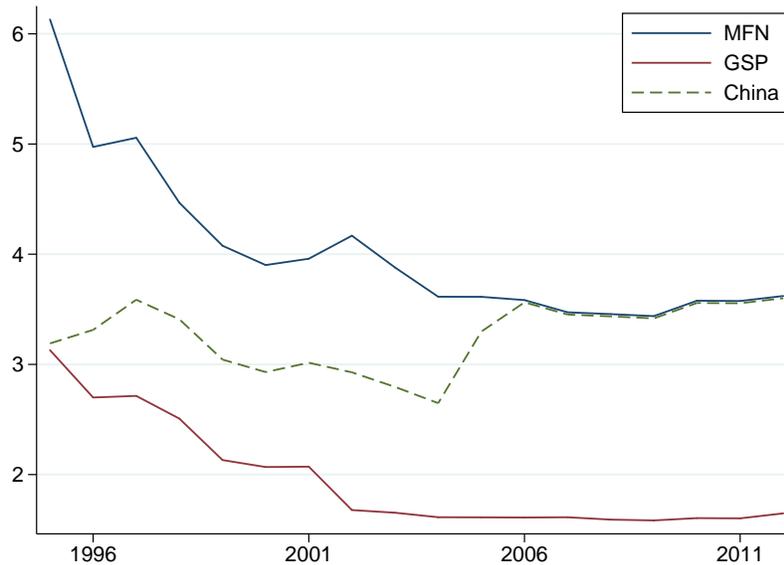
an average manufacturing good. The former gradually decline and evolve proportionally.[5] Tariffs on Chinese goods declined at a lower rate until 2004. Since 2005 they increased due to graduation from GSP preferences and have fully returned to MFN rates since 2006. Altogether, the European trade environment, in terms of applied tariffs, appears to have worsened for China, and so it can barely explain China's rising exports to the EU since 2002.

3   Policy Spillover

This section attempts to rationalize the rise of Chinese exports to the EU. It presents first the bilateral framework similar to Handley and Limão (2013) and then expands on it to establish a channel through which bilateral policies affect the multilateral export performance.

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[5]The local MFN peak in 2002 is driven by a number of steel products where the EU temporarily raised the tariff rate. This occurred in response to the rise of US steel tariffs in March 2002 under president George W. Bush.

3.1 Baseline Model

3.1.1 Setup

Demand. Following the Melitz (2003) framework 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. Monopolistic firms produce only one variety and charge a mark-up over their marginal costs in order to maximize profits. 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.

Firm Entry. The profit function of the firm is $\pi = (\tilde{p} - c)x - f$. The unit cost parameter is given by $c_{jn} \equiv d_{Jn} \tau_{Jn}(w/\varphi_j)$ whereas the unit price $\tilde{p}_{jn} \equiv p_{jn}/\tau_{Jn}$ received by the firm is discounted by the tariff collected at the border. Substituting (2) and (3) into the profit

function states the problem of the firm that considers exporting to n

$$\pi_{jn} = \tau_{Jn}^{-\sigma} \left(\frac{d_{Jn}}{\varphi_j}\right)^{1-\sigma} E_{Jn} (1-\epsilon) \left(\frac{w}{P_{Jn}\epsilon}\right)^{1-\sigma} - f_{Jn}. \tag{4}$$

A positive fixed cost $f_{Jn} > 0$ prevents firms from exporting to n when operating profits are too low. Using $\pi_{jn} = 0$ identifies the marginal firm which is indifferent between exporting and not-exporting. It has productivity

$$\varphi_{Jn}^* = \tau_{Jn}^{\frac{\sigma}{\sigma-1}} \left[\frac{f_{Jn}}{E_{Jn}(1-\epsilon)}\right]^{\frac{1}{\sigma-1}} \left(\frac{d_{Jn}w}{P_{Jn}\epsilon}\right). \tag{5}$$

Higher applied tariffs, τ, shipping costs, d, or fixed costs, f, require a higher firm productivity to pass the zero-profit cutoff (ZPC). Higher demand E or prices P in the destination market allow less productive firms to export profitably.

3.1.2   Tariff Uncertainty

The analysis of tariff uncertainty considers the possibility that τ^s takes different values depending on the policy regime faced in destination n: $s = \{p, np\}$. If the importing country grants preferential 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 focusses on the removal of uncertainty regarding the application of preferential tariffs. As long as this uncertainty exists, firms do not know how future tariffs will be. They assume that a shift from preferential to non-preferential tariffs occurs with probability $0 \leq \delta \leq 1$. The expected tariff can be written as a weighted geometric average of the two scenarios, $\tau^E = (\tau^{np})^\delta (\tau^p)^{(1-\delta)}$. This assumption is made to simplify the exposition, as well as the application of the model to the data.[6] It implies that

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[6]The characterization of uncertainty in related papers is more explicit. Handley and Limão (2013) analyze alternative policy regimes and their probabilities within a Markov transition matrix, but they rule out by assumption several transitional trajectories. Feng et al. (2014) model tariff uncertainty as a Poisson-process with an arrival rate λ. In their empirical applications both papers compare scenarios analogous to moving from $\delta > 0$ to $\delta = 0$, so that the simplifying notation used here should be sufficient to illustrate the policy

$\tau^p \leq \tau^E \leq \tau^{np}$. Equation (5) can be rewritten as

$$\varphi^*_{Jn} = (\tau^E_{Jn})^{\frac{\sigma}{\sigma-1}} \left[\frac{f_{Jn}}{E_{Jn}(1-\epsilon)} \right]^{\frac{1}{\sigma-1}} \left(\frac{d_{Jn}w}{P_{Jn}\epsilon} \right) \tag{6}$$

and gives the key result of the bilateral model with tariff uncertainty.

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

This follows from the positive relationship between (expected) tariffs and the productivity threshold, $(\partial\varphi^*/\partial\tau^E) > 0$, and from the fact that expected tariffs equal the preferential rate when uncertainty vanishes ($\delta = 0$). If preferential and non-preferential tariffs are the same, uncertainty has no effect on the productivity threshold, and nothing happens when it is removed.

3.1.3 Product-level Predictions for Bilateral Trade

Firm-level export revenues $r_{jn} \equiv \tilde{p}_{jn}(\varphi)x_{jn}$ can be aggregated to obtain product-level predictions

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

where $a_{Jn} \equiv \tau^{-\sigma}_{Jn} d^{1-\sigma}_{Jn} A_{Jn}$ and $A_{Jn} \equiv \left[\left(\frac{\sigma}{\sigma-1} \right) \frac{w}{P_{Jn}} \right]^{1-\sigma} \frac{E_{Jn}}{\sigma}$. The expression in parentheses is equivalent to multiplying the total number of firms, 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 productivities levels across firms are Pareto distributed,[7] the probability that a random productivity draw from this distribution exceeds its lower bound φ_L equals $G(\varphi) = \left(\frac{\varphi_L}{\varphi} \right)^k$.[8]

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spillover in the next subsection.

[7]This is a standard assumption in the context of generating aggregate predictions from the Melitz (2003) model, and it is valid at least for the right tail of the distribution, where exporting firms typically reside.

[8]This feature suggests that the probability of a firm to have a certain productivity level $\varphi_j > \varphi_L$ decreases with the size of φ_j/φ_L. The shape parameter k raises this ratio to a power $k > (\sigma - 1)$ so that a larger k

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. A gravity equation is obtained by plugging (6) into (9) and taking logs

$$\ln R_{Jn} = -\frac{\sigma k}{\sigma-1} \ln \tau_{Jn}^p - \delta_n \frac{\sigma k}{\sigma-1} (\ln \tau_{Jn}^{np} - \ln \tau_{Jn}^p) - k \ln d_{Jn} + \frac{k}{\sigma-1} \ln A_{Jn} + \ln \alpha_J - \frac{k-\sigma+1}{\sigma-1} \ln f_{Jn}. \tag{10}$$

Equation (10) illustrates how the removal of tariff uncertainty in destination n affects exports to that country. As $\delta_n \to 0$, the second term of the right-hand side of the equation disappears, and exports to n increase. The log-difference between non-preferential and preferential tariffs measures the "tariff threat" exporters face under uncertainty. Overall, the removal of tariff uncertainty reduces the expected tariff rate. Accordingly, also the ZPC goes down which implies that the adjustment takes place at the extensive margin. These predictions have been confirmed for China's exports to the US in previous studies (Handley and Limão, 2013; Feng et al., 2014).

3.2 Separable Fixed Costs and Multilateral Trade

The baseline model provides no explanation for increased exports to a country where policies did not change. To establish this link, additional structure is imposed on the fixed market-entry costs f_{Jn}. It is assumed that f_{Jn} can be separated into a local and a global component, $f_{Jn} \equiv f_n + f_J$.[9] With this assumption the export decision of a firm becomes interdependent. The global fixed cost component has to be paid irrespective of the number of destinations a firm serves. It implies that the burden of the global fixed cost can be distributed across

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implies a higher concentration of low-productivity firms and a smaller number of very productive firms.

[9]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.

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 the global fixed cost:

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

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

$$
\Phi_N^* = \sigma^{\frac{1}{\sigma-1}} \left(\frac{w}{\epsilon}\right) \left[\sum_{n=1}^{N} \tau_{Jn}^{\frac{\sigma}{\sigma-1}} \left(\frac{f_n + f_J}{E_{Jn}}\right)^{\frac{1}{\sigma-1}} \frac{d_{Jn}}{P_{Jn}}\right].
\tag{12}
$$

3.2.1   General Implications of the Multilateral Productivity Threshold

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

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). A row with a negative partial profit lowers total exporting profits. It implies that the number of destinations an exporting firm serves results from an assessment of each market. To determine this number, destinations can be ranked in decreasing order of the bilateral partial profits. The 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 4: Partial and Total Export Profits of Two Firms
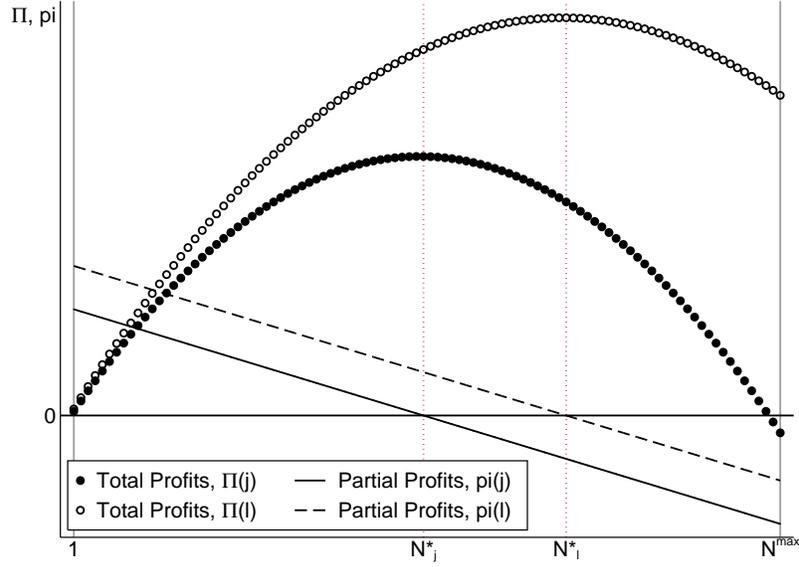
Note: Author's calculations based on Eq. (11) with 100 destinations.

Figure 4 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 its total export profits earned by exporting to a respective number of destinations. Total profits rise, up to the point where partial profits become negative. The firm's optimal number of destinations is indicated at point N_j^*. The other firm l has productivity $\varphi_l > \varphi_j$ and serves a larger number of countries. This prediction conforms to evidence from firm-level data (e.g. Eaton et al., 2004). It leads to a statement on the multilateral productivity threshold in Equation (12):

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

Lemmas 2-4 hold also for the case where fixed costs are purely country-specific. The distinct feature with a global fixed cost component arises when f_J is large enough to prevent a firm from earning positive profits in the first destination $\tilde{\pi}_{j1} - f_J < 0$. If that is the case, a firm

might need the revenues from several markets in order to export profitably. This prediction seems to be at odds with observations where firms export first to one destination and enter additional markets when they grow. However, it is possible that Chinese manufacturers are differnt. Amiti and Freund (2010) note that a large fraction of China's recent export growth is driven by processing trade. The firms carry out certain production processes but contribute less to the development of a product or to distributing it to final customers. This allows them to save destination specific costs by focussing on standardized processes. The global fixed cost f_J is paid to attract orders from foreign firms that are willing to save labor costs through outsourcing.[10] The bilateral fixed cost f_n governs the eligibility of Chinese firms to be integrated in the production chain of certain countries.

3.2.2 Bilateral Tariff Uncertainty and Multilateral Trade

The effect of a reduction in tariff uncertainty on multilateral trade is illustrated with two examples. The first assumes that countries are symmetric. The second example considers the market size of the policy making country.

Two Symmetric Countries.   Supposing that firms consider exporting to two foreign destinations, $n = \{1, 2\}$, the baseline scenario describes the outcome where the tariffs are uncertain in country 1. The applied tariff in country 1 corresponds to the preferential rate $\tau_1^p = 1$ but the non-preferential rate $\tau_1^{np} = 2$ might be applied with a probability $\delta_1 = 0.5$. The expected tariff in country 1 is 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 the ZPC productivity for any possible scenario. To obtain numerical results it is assumed that the elasticity of substitution is $\sigma = 3$ and that all other non-tariff

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[10]Empirical evidence along these lines is provided for Mexican firms that make investments to improve the quality of their products before they start to export (Iacovone and Javorcik, 2012). Feng et al. (2012) find that Chinese exporters benefited from importing larger amount of intermediate inputs after they entered the WTO.

variables equal 1.[11]

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.

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 most firms would export only to country 2 because column (3) states the lowest threshold. The second row displays the ZPCs when uncertainty is removed. It indicates that most firms will export to both destinations, as shown in column (1). The regime shift from uncertainty to certainty in country 1 induces firm entry into both markets. This follows because the ZPC without uncertainty is below any other ZPC of the baseline. The firms that become exporters after the policy change have productivity $\varphi \in (2.90; 3.18)$.

Three Asymmetric Countries. 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 5 illustrates how the removal of tariff uncertainty translates to multilateral exports. It depicts the (log of the) Pareto-density function $g(\varphi)$. Typically, most firms have a relatively low level of productivity and very few firms are very productive. The baseline scenario is

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[11] $\sigma = 3$ 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).

Figure 5: Multilateral Productivity Thresholds with Bilateral Tariff Uncertainty and Asymmetric Countries
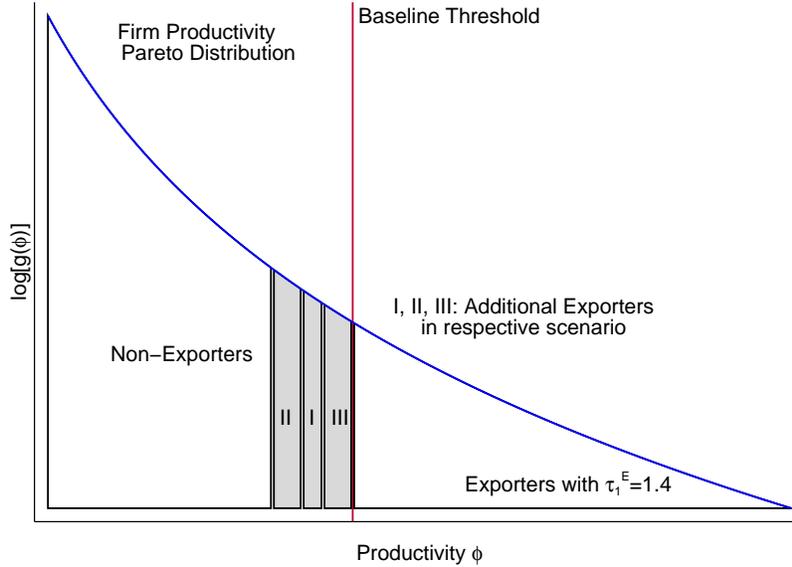
Note: Author's calculations based on Eq. (12) with three asymmetric destinations (see text). Figure shows Pareto density function $g(\varphi) = \frac{k}{\varphi_L} \left(\frac{\varphi_L}{\varphi}\right)^{k+1}$ with lower bound $\varphi_L = 1$ and shape-parameter $k = 2$. The vertical axis is scaled in logs.

normalized and shows the ZPC applicable for exporting to countries 1 and 2.[12] 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 small. The largest amount of new exporters is found for case II in which country 1 is big.

The figure suggests that the size of a market is correlated with the fraction of global fixed costs it absorbs. If the policy-making country is very small, the effect of removing TPU might be negligible. Scenario III suggests that uncertainty in country 1 induces 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. As a result, exports do not adjust at the multilateral level, because there is no additional entry of firms

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

that were non-exporters in the first place.

The results for asymmetric countries in terms of market size carry over to heterogeneity in terms of other country characteristics (e.g. bilateral trade costs). To evaluate the predictions of the model, two hypotheses shall be spelled out:

Proposition 1 *If $\tau_{Jn}^p < \tau_{Jn}^{np}$, a removal of tariff uncertainty in a large country n induces an increase in the value of exports to any other destination $m \neq n$ with positive partial profits.*

Proposition 2 *If $\tau_{Jn}^p < \tau_{Jn}^{np}$, a removal of tariff uncertainty in a large country n induces an increase in the number of destinations N to which product J is exported.*

Both propositions follow from a new ranking of a firm's bilateral partial profits and the possibility of re-distributing global fixed costs. As tariff uncertainty vanishes, non-exporters are able to take into account additional revenues, enabling them to earn positive total exporting profits. The policy spillover works through a reduction of the multilateral export threshold. This implies that the adjustment takes place at the extensive margin. An investigation of the removal of US tariff uncertainty towards China will provide evidence on the existence of countries that are large enough to generate the predicted effects.

4 Empirical Framework

4.1 Empirical Model

The empirical specification is derived from the previous section. It extends the gravity equation stated in Equation (10) by the global fixed cost.

$$\ln R_{Jn} = -\frac{\sigma k}{\sigma - 1} \ln \tau_{Jn}^E - k \ln d_{Jn} + \frac{k}{\sigma - 1} \ln A_n + \ln \alpha_J - \frac{k - \sigma + 1}{\sigma - 1} \ln \left(f_n + \theta_{Jn} f_J \right) \quad (13)$$

The parameter θ_{Jn} indicates that a fraction of f_J is allocated to each destination n. It implies $0 < \theta < 1$ for any existing trade relationship and $\theta_{Jn} = 0$ whenever $\tilde{\pi}_{Jn} \leq 0$.

An allocation rule could take the following form:

$$\theta_{Jn} \equiv \frac{\tilde{\pi}_{Jn}}{\Pi_J - f_J},$$

which ensures that the rankings of bilateral partial profits $\tilde{\pi}_{Jn}$ and of the same profits net of global costs, $\tilde{\pi}_{Jn} - \theta_{Jn} f_J$, are proportional. The interpretation is that a reallocation of the fixed cost burden away from destination n increases exports to this country $\partial R_{Jn}/\partial \theta_{Jn} < 0$. To incorporate the effect of θ into the estimation equation, it is assumed that θ_{Jn} is correlated with the tariff threat faced in a country $m \neq n$. In the context of this paper: $\partial \theta_{Jn}/\partial GAP_J^{US} > 0$. The tariff uncertainty Chinese exporters faced in the US forced them to increase θ for any other export destination.

The econometric analysis focusses on the removal of uncertainty. Hence, the variable GAP_J^{US} is interacted with a period dummy D_t^T which equals zero before China's WTO entry and one afterwards. This is shown by the first term of the following equation:

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

The sign of b_1 is expected to be positive, since the removal of US tariff uncertainty reduces θ_{Jn} and increases R_{Jn}. The second term meausures the tariff rate in destination n and should reveal a negative effect. The analysis focusses on China's exports to EU countries, so it is assumed that there is no tariff uncertainty, and $\ln \tau_{Jn}^E = \ln \tau_{Jn}^p$. The product-destination fixed effect captures the country specific trade costs of shipping good J to its destination; $b_{Jn} \equiv -k \ln d_{Jn}$. Destination-specific time effects $b_{nt} \equiv \frac{k}{\sigma-1} \ln A_{nt}$ control for changing aggregate conditions in each destination, such as higher demand or prices. b_{Jn} and b_{nt} absorb also the local fixed cost component, f_n. Finally, sector-time effects b_{St} capture variation in α_J over time at a more aggregated level.

4.2 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}, on applied EU tariffs for China, and on the value of Chinese exports to the EU.

US tariff data is drawn from the National Bureau of Economic Research's (NBER) website (Feenstra et al., 2002). The same data was used by Pierce and Schott (2013) to study the effect of removing US tariff uncertainty on domestic manufacturing employment. The data reports ad-valorem equivalents of applied tariffs at the 8-digit level of the Harmonized Tariff Schedule (HTS). This allows matching the data up to the 6th digit with product codes of the Harmonized System nomenclature (HS6). The US tariff threat is calculated at the disaggregated level and then averaged over the respective HS6 category J:

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

The definition $GAP_J^{US} \equiv \ln \tau_J^{Col2} - \ln \tau_J^{MFN}$ follows from the definition of the expected tariff rate, $\ln \tau^E = \ln \tau^p + \delta(\ln \tau^{np} - \tau^p)$, where the expression in parentheses equals GAP. Following other studies, the year 1999 is chosen to compute US tariff uncertainty. It varies only across product categories J but will be interacted with the period dummy for China's WTO membership. In a final step the HS6 codes are harmonized over time to reflect the classification of goods according to the HS 1988/1992 revision.[13]

European tariff data is obtained from the tariff schedules at the World Integrated Trade Solution (WITS). It reports ad-valorem equivalents at the 8-digit level of the Combined Nomenclature (CN8). Also this classification can be matched with HS6 products at the first six digits. A problem with this data is that information is missing for some years and products, and that this is difficult to trace, because the raw data does not take into account the revisions of product codes over time. Using the correspondence tables provided

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[13]Correspondence tables for the HS nomenclature are available at the United Nations Statistics Division (UNSD).

by Bernard et al. (2012) helps to trace most products.[14] A few remaining empty cells could be filled using information from similar products where observable tariffs evolved identically. The complete set of MFN rates is needed to compute GSP rates and those applicable to China in each product year. The GSP discount factors and the exceptions for China were drawn from the respective European Council Regulations. After aggregating and harmonizing product codes according to the HS6 1988/1992 revision, tariffs were expressed as $\tau = 1+(\%\text{-rate}/100)$.

The data on Chinese exports was obtained from the UN Comtrade Online Database. Exports to the 15 EU members were selected and product codes were converted into the HS6 1988/1992 revision. Belgium and Luxembourg are treated as one country because they are not reported separately in all years. To distinguish the extensive margin, missing information on Chinese exports for a given HS6-country pair is assumed to reflect that no trade had taken place. In a fully balanced panel, covering the years 1995-2005, observations with zero trade amounts to 56 percent. Most export zeros are observed for Ireland (75%) and Austria (71%) and the fewest are reported for exports to Germany (34%) and Italy (40%). These numbers generally decrease over time.

5 Results

5.1 Main Findings

5.1.1 Level of Chinese Exports to the EU

Table 2 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 $3,985$, comprised in the HS Chapters 28-96. About six percent of the products were not exposed to US tariff uncertainty. The tariff threat for an average exposed product was 0.289 in 1999. The reported results represent a sample with 14 destinations n and 11 sectors S.

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[14]The CN8 correspondence tables for the most recent years are available at the EUROSTAT Reference and Management of Nomenclatures (RAMON) archive.

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 18.7 percent relative to a non-threatened product. Potentially differential patterns may arise in the textile and clothing sector (T&C) where the EU removed 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). The estimated effect is lower but still statistically significant and positive. Column (3) controls explicitly for the removal of quotas.[15] It suggests that China's exports increased due to both the removal of quotas and the removal of US tariff uncertainty. The implied average increase is 11.7 percent relative to non-threatened products.

The estimated coefficient can be interpreted in the context of the theoretical model. According to Equation (13) it reflects the elasticity of exports with respect to a reduction of the fixed-costs burden; $(k - \sigma + 1)/(\sigma - 1)$. Column (3) suggests that a one percent reduction increases exports by 0.405 percent. Head et al. (2014) estimate the Pareto parameter of China's firm productivity distribution to be $k = 4.854$. The implied elasticity of substitution is $\hat{\sigma} = 4.455$; not too far from to their parametrization ($\sigma = 4$).

Columns (4) through (6) adopt a discrete measure of the US tariff threat. Similar to Handley and Limão (2013) GAP is divided into groups. The first group (not shown) considers the goods where $GAP = 0$, i.e. where Column 2 rates equalled the MFN tariff rate. The second group considers the bottom quartile and the last group includes the top quartile of the tariff threat. Half of the products fall into the third group, where GAP ranges between 0.19 and 0.35. All specifications confirm the qualitative results of the baseline specification. Quantitatively, they suggest that US tariff uncertainty prevented market entry especially when $GAP \geq p[25]$. This corresponds to a difference of 0.187 log points or more between Column 2 and MFN tariffs. For an average good in the top quartile, column (6) suggests that exports increased by 17.3 percent relative to the non-threatened products. Table 2 also

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

Table 2: Chinese Exports to the EU-15 after the Removal of US Tariff Uncertainty; Linear Panel Regression, 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	0.647**	0.237**	0.405**			
	(0.076)	(0.091)	(0.077)			
$p[+]$-$p[25]$				0.192**	0.158*	0.212**
				(0.074)	(0.076)	(0.073)
$p[25]$-$p[75]$				0.400**	0.371**	0.387**
				(0.071)	(0.073)	(0.071)
$p[75]$-$p[100]$				0.463**	0.311**	0.393**
				(0.072)	(0.076)	(0.072)
EU Tariff	-0.386	0.283	-0.052	-0.524	0.088	-0.181
	(0.420)	(0.433)	(0.419)	(0.421)	(0.435)	(0.420)
EU Quota '02			0.576**			0.576**
			(0.034)			(0.034)
EU Quota '05			0.449**			0.446**
			(0.050)			(0.049)
Observations	270,767	207,476	270,767	268,499	205,966	268,499
R-squared	0.170	0.176	0.172	0.171	0.177	0.173
Fixed effects	All	All	All	All	All	All

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. Robust standard errors in parentheses; significance: [a] $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

suggest that applied tariffs in the EU had no effect on China's exports to the EU. This is not surprising given that they changed only marginally during this period.

5.1.2 Extensive vs. Intensive Margin

To analyze the trade creation effect of the US policy change, patterns at the extensive margin are investigated. Table 3 presents results of different specifications that capture this adjustment.

The first two columns present the odds-ratio and the coefficient of a logistic regression. The dependent variable takes a value equal to one when China exports to a given product destination at time t, and zero otherwise. Column (1) suggests that the removal of non-tariff trade barriers, e.g. US tariff uncertainty and EU import quotas, increase the probability of

Table 3: Chinese Market Entry in the EU-15 after the Removal of US Tariff Uncertainty; Alternative Estimators, 1995-2005

| | Logistic Regressions | | Linear Regressions | | |
| | Odd Ratio | Coeff. | # Destinations | Norm. Growth | Log Growth |
	(1)	(2)	(3)	(4)	(5)
US Tariff Threat	2.511**	0.921**	1.770**	0.159**	0.094[a]
	(0.201)	(0.080)	(0.275)	(0.034)	(0.048)
EU Tariff	0.406*	-0.902*	0.300	-0.010	-0.327
	(0.169)	(0.417)	(1.071)	(0.327)	(0.405)
EU Quota '02	1.157*	0.146**	0.920**	0.090**	0.082**
	(0.040)	(0.034)	(0.144)	(0.017)	(0.023)
EU Quota '05	2.115**	0.749**	1.234**	0.691**	0.683**
	(0.148)	(0.070)	(0.210)	(0.039)	(0.060)
Observations	341,814		44,038	284,134	204,837
R-squared	0.177		0.364	0.056	0.010
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 results of a logistic regression. Column (3) considers the number of EU-15 destinations served by China. The last two columns compare the normalized (4) and the log-growth rate (5) of Chinese to the EU-15 over time. Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year. Standard errors in parentheses; significance: [a] $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

observing Chinese exports to a given European product-country pair. Higher tariffs make trade less likely to occur. The signs of the coefficients stated in column (2) show the marginal effects of a change in the independent variable on the odds ratio. As expected, tariffs reduce the probability of exporting to the EU, whereas the removal of other trade barriers reveals positive and statistically significant effects. Because the logistic regressions were not able to include a complex fixed effects structure, column (3) considers an alternative specification. The dependent variable corresponds to the number of EU destinations to which China exports good J. This reduces the dimension of the panel to have variation across products over time. The result confirms that products with a high US tariff threat before China's WTO entry were exported to more destinations after the removal of the threat. The same is found for the removal of import quotas on T&C products. 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. Column (5) uses

the log-difference as a measure of the growth rate. By construction, it is confined to report patterns at the intensive margin, because it cannot be computed when trade was zero at the beginning or at the end of the period under consideration. The effect of the US policy change is only marginally significant in statistical terms. In contrast, column (4) uses the normalized growth rate (Davis et al., 1998; Pierce and Schott, 2013). It takes into account entry and exit in export markets by taking its bound values $g^N \in [-2, 2]$. The rate is defined as

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

If this growth rate is considered, the removal of US tariff uncertainty reveals a positive and statistically significant effect. This result confirms that the US policy change encouraged Chinese exporters to establish new trade relationships with EU countries.

5.1.3 Robustness

The removal of US tariff uncertainty was not the only change in trade policies that could have caused China's export boom after WTO accession. A number of alternative explanations are taken into consideration to investigate whether the policy spillover picks up the effect of other changes in China's export oriented industries. Details on the data sources and construction of variables are provided in the appendix to this paper.

Unobserved Productivity Effects upon WTO Accession. The first test analyzes whether China's WTO accession has had any productivity effects due to a general reduction of economic uncertainty. Such a general effect might reveal structural patterns (e.g. due to comparative advantage) and should be visible also in relative export prices (e.g. due to investments into better technologies). Table A1 shows the results from including an interaction of China's *inverted* revealed comparative advantage (RCA) with the indicator variable for China's WTO membership. The inverted measure was chosen to point out the substan-

tial structural transition undergoing China's exports since the early 1990s.[16] While exports of comparative disadvantage goods, as measured for the years 1995-1997, grew significantly faster after China's WTO accession, the removal of US tariff uncertainty continues to be statistically significant. In columns (4) through (6) of the table, it can be observed that a depreciation of China's currency towards the currencies of the European importers has contributed to this transition, facilitating export growth in comparative disadvantage industries.

A similar result is obtained when changes in Chinese relative export prices are taken into consideration (Table A2). The data uses quality-adjusted export prices from Feenstra and Romalis (2014). Their change between 1997-1999 and 2002-2004 was interacted with the WTO dummy. The results shown in columns (1) through (3) show how decreasing export prices facilitated exports to the EU. In columns (4) to (6) changes in EU countries' relative import prices are used to deflate export price changes, and the pattern can be confirmed. In all specifications, statistical significance and estimated magnitude of the effect of the US policy change remain unaffected.

Removal of Investment Restrictions in China. An alternative explanation for the rise of Chinese exports could be changes in the business environment in China. In particular, upon WTO entry, China complied to an equal treatment of foreign and domestic investors. This may have encouraged offshoring and other investments into China's manufacturing industries. I follow Pierce and Schott (2013) by approximating products' elasticity to investment restrictions using measures of products' contract intensity from Nunn (2007), and interacting them with the WTO dummy. It should be expected that investment and exports rise more for products where the contract intensity is high.

Table A3 shows the results from re-estimating (14), augmented by the interaction of products' contract intensity with China's WTO dummy. Columns (1) through (3) of the table refer to the strict measure of contract intensity, which measures the share of differentiated

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[16]At a more aggregate level, it can be seen that China's share of Textile, Clothing, and Apparel in total manufacturing exports declined from 45 to about 20%, while the share of machinery and electronics products grew up to 40%, by 2012.

input goods used in the production. Columns (4) through (6) consider the broader measure, which is based on the cumulated share of differentiated and reference-priced input goods. Contract intensity is statistically significant, suggesting that the rise of Chinese exports to the EU is partly driven also by China's removal of foreign investment restrictions. The table confirms the statistical validity of the US policy spillover, except for one specification, in column (5), where the effect is only marginal. However, since the broader measure includes inputs with an intermediate degree of diversification, and hence contracting intensity, is not entirely clear whether the result unambiguously captures the removal of Chinese investment restrictions.[17]

Export Subsidies. Production subsidies may have led to higher exports in some product categories. To investigate such effects, data from Bown (2014) and Bown (2015) are used to control for HS6 products subject to anti-dumping (AD) and countervailing duty (CVD) measures in the European Union. AD-filings of products exported from China to the EU could be observed throughout the 1990s and 2000s. Columns (1) through (3) of Table A4 indicate that exports of these goods decreased while they were subject to investigations, until the file was closed or countermeasures were revoked. In columns (4) through (6), additional information is included, for goods which became subject to CVD, since 2010. Although the data for European CVD cases starts only after the first decade of the 2000s, interacting these products with the WTO indicator suggest that China may have subsidized their exports starting from an earlier time.

In all specifications, the estimates for the US policy change remain positive and statistically significant, and in similar orders of magnitude than in the baseline estimations. In fact, the explanatory power of the empirical model remains merely unchanged. This might indicate that the indirect information on export subsidies does not cover all of the relevant products.

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[17]Moreover, estimating the same equation based on a restricted sample suggests a robust US policy spillover at the 5 percent level or better (see Appendix for details).

5.2 Further Results

The main findings could not reject either proposition made at the end of Section 3. This subsection attempts to provide further results that support the mechanism emphasized throughout this paper.

5.2.1 US Share in Chinese Exports

A main concern could be that the US policy change operates through a different mechanism than the redistribution of a global fixed costs burden. The problem is that such costs cannot be observed. To provide further (suggestive) evidence, the proposed allocation rule for θ_{Jn} will be addressed. The interpretation was that the US policy change invoked the allocation of a larger fraction of f_J to the US so that EU markets become easier to penetrate. This fraction is assumed to be proportional to the share of the partial bilateral profit in total export profits. To analyze this, the explanatory variable GAP_J^{US} is replaced by the fraction of China's total exports of product J that is shipped to the US before and after WTO entry.

The variable is constructed by calculating the average $s_{J,US} = R_{J,US}/R_J \approx \theta_{J,US}$, respectively for two periods 1992-2000 and 2002-2009. For the first period it yields $\bar{s}_{J,US}^{pre}$, the fraction of exports of J shipped to the US before China's WTO entry. For the second period $\bar{s}_{J,US}^{post}$ denotes the fraction of exports of J shipped to the US after China's WTO entry. In order to capture the change of this fraction the difference $\Delta \bar{s}_J^{US} = \bar{s}_{J,US}^{post} - \bar{s}_{J,US}^{pre}$ is used to replace GAP_J in Equation (14). A positive $\Delta \bar{s}_J^{US}$ suggests that China exports more of good J to the US compared to the years before its WTO membership. This implies that a larger fraction of the global fixed cost burden is covered through this trade relationship.

The results are presented in Table 4. Column (1) is analogous to the third column of Table 2. The estimated coefficients for EU tariffs and the quota removals are qualitatively and quantitatively unchanged. The coefficient obtained for the US share in Chinese exports is positive and statistically significant. Moreover, it exceeds that obtained from the original specification. An explanation could be that $\Delta \bar{s}_J^{US}$ can take negative values allowing it to

Table 4: Chinese Exports to the EU-15 and the Role of Trade with 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	0.789**	3.371**	1.215**	2.073**	0.030	0.039
	(0.090)	(0.293)	(0.087)	(0.348)	(0.037)	(0.058)
EU Tariff	-0.073	0.429*	-0.847*	0.403	-0.000	-0.328
	(0.420)	(0.180)	(0.419)	(1.068)	(0.327)	(0.406)
EU Quota '02	0.595**	1.222**	0.201**	1.029**	0.106**	0.091**
	(0.034)	(0.040)	(0.033)	(0.142)	(0.016)	(0.023)
EU Quota '05	0.448**	2.102**	0.743**	1.228**	0.694**	0.685**
	(0.050)	(0.147)	(0.070)	(0.213)	(0.039)	(0.060)
Observations	267,870	337,711		43,307	281,203	202,702
R-squared	0.173	0.183		0.374	0.056	0.010
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] $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

better describe the conditions of Chinese exporters willing to export to the EU. Related to this, the removal of US tariff uncertainty does not explain the entire pattern of Chinese exports to the US, so that $\Delta \bar{s}_J^{US}$ provides a more complete assessment of the theoretical model. The remaining columns are analogous to Table 3. Except for columns (5) and (6) the previous results are confirmed. Together this analysis suggests that Chinese exporters benefit from selling much of their exports in the US. This supports the hypothesis of an existing global fixed cost burden and the associated economies of scale.

5.2.2 Dynamic Adjustments and Transitional Growth

So far, the analysis has concentrated on the years immediately after China's WTO accession. Available trade and tariff data allows an extension of the analysis up to the year 2012 so that the effect can be studied over a longer period. 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 $\hat{b}_1 = 0.691$ (see Table 5). It is larger and statistically different from

that obtained for the baseline period shown in Column (3) of Table 2. One explanation could be that the US policy change interacts with a dynamic component. While Section 3 considers comparative statics, it is possible that firms gradually increase their production capacities and exploit their full cost advantage with a delay. A complementary explanation would be that some firms start to export only when they observe the success of other firms (e.g. Hausmann and Rodrik, 2003). Following these arguments, a dynamic specification should eliminate the positive correlation between estimates of b_1 and the length of the post WTO-entry period.

Including a lagged endogenous variable $\ln R_{Jnt-1}$ on the right-hand side of the estimation equation may create problems due to correlated errors which lead to biased coefficients. However, biased coefficients can be controlled by inferring the upper and lower bounds of the true coefficient of $\ln R_{Jnt-1}$. It should lie between the estimate obtained from a dynamic pooled OLS model (POLS) and a dynamic fixed effects (FE) model (Roodman, 2009). The former will overestimate the true coefficient, while the latter will produce a downward biased estimate for the lagged endogenous variable. Hence, a downward-biased dynamic specification will pick up some but not all of the dynamics that were induced by the US policy change. It should produce more consistent coeffients $\hat{b}_1$ as the period under study is extended to more recent years.

The left panel of Table 5 shows the results for the full 1995-2012 period, for each of the three specifications considered: baseline in column (1), dynamic FE in column (2), and dynamic POLS in column (3). The POLS model generates a higher coefficient for the lagged endogenous variable than the FE model. In the two dynamic specifications, the removal of the US tariff threat reveals a lower coefficient for GAP than in the baseline. However, the POLS model produces unplausible results for the effect of EU tariffs, and also the effects of the removal of quotas in the EU become fragile. It seems that the FE model in column (2) produces more plausible results overall.[18]

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[18]Assuming that the baseline specification in column (1) represents the long-run effect of the policy spillover, the estimate must be compared to $b_1^{long} = b_1/(1-\gamma)$, where γ denotes the coefficient for the lagged endogenous variable. For column (2), the implied long-run effect is 0.543. This indicates that the coefficient in column (1) is biased upwards.

Table 5: 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	$T = 2002$	$T = 2006$	$T = 2010$
	(1)	(2)	(3)	(4)	(5)	(6)
US Tariff Threat	0.691**	0.303**	0.231**	0.336**	0.120**	0.030
	(0.085)	(0.056)	(0.023)	(0.072)	(0.031)	(0.026)
EU Tariff	-3.050**	-2.024**	0.867**	-0.056	-0.205	-0.005
	(0.420)	(0.312)	(0.110)	(0.397)	(0.319)	(0.270)
EU Quota '02	0.391**	0.176**	0.017	0.287**	0.068**	0.068**
	(0.038)	(0.026)	(0.013)	(0.035)	(0.016)	(0.014)
EU Quota '05	0.484**	0.352**	0.035*		0.396**	0.241**
	(0.046)	(0.032)	(0.017)		(0.026)	(0.016)
EU Quota '09	0.870**	0.514**	0.188**			0.046
	(0.071)	(0.041)	(0.028)			(0.029)
Lagged Exports		0.442**	0.825**			
		(0.003)	(0.001)			
Observations	508,433	414,950	414,154	182,946	321,436	472,600
R-squared	0.313	0.474	0.746	0.045	0.078	0.116
Fixed effects	Jn, nt, St	Jn, nt, St	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] $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

Figure 6 presents the point estimates of the effect of GAP_J^{US} graphically. In Panel (a) it shows results analogous to columns (1) and (2) of Table 5 where the sample period was expanded stepwise ($T = 2002, ..., 2012$). The dashed line shows how the baseline specification produces higher coefficients $\hat{b}_1$ as the period gets longer. The immediate effect for $T = 2002$ is $\hat{b}_1^{02} = 0.214$. In the dynamic FE model the immediate effect is the same and remains within the confidence intervals of any other period length. This is indicated by the solid line in Panel (a) where $\hat{b}_1$ still increases over time but remains statistically the same. Together with $k = 4.854$ $\hat{b}_1^{02} = 0.214$ implies that $\hat{\sigma} \approx 5$. This estimate is still plausible given that estimates vary considerably across countries (Imbs and Méjean, 2010).

The right panels of Table 5 and Figure 6 analyze the evolution of the growth rate. As

Figure 6: Evolution of the US Policy Effect on Chinese Exports to the EU; 2002-2012

(a) Level Effect
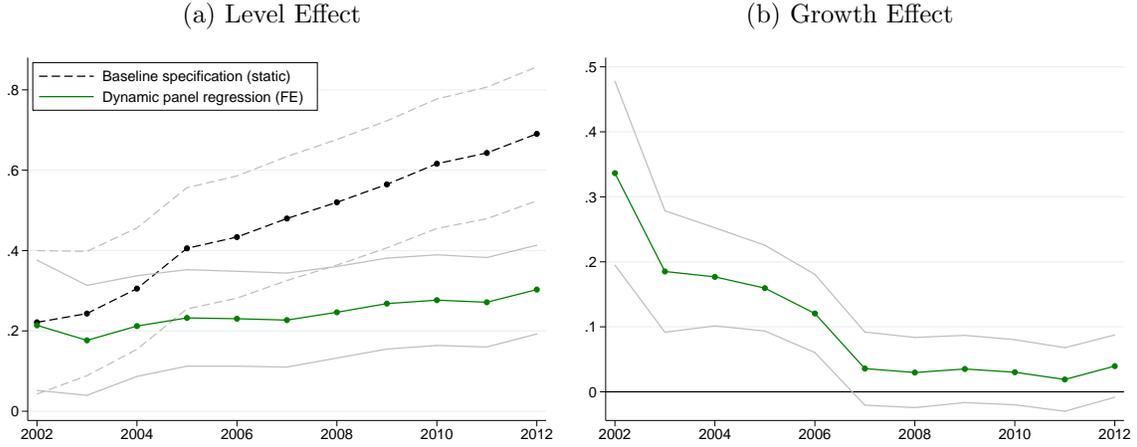
(b) Growth Effect

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.

was shown in column (4) of Table 3, the average annual growth rate of products exposed to high uncertainty increased after China's WTO entry. If this effect was due to the policy change, it should disappear as the year of the WTO accession moves further into the past. Columns (3) to (6) show econometric results for the normalized growth rate. In the first year of China's WTO membership, the estimated coefficient is between 2.5 and 3 times larger than that of the years up to 2006. The last column shows that growth rates do not differ when the period is extended up until 2010. Panel (b) of Figure 6 shows how the effect phases out and disappears after 2006. Together, the results presented in this subsection suggest that the analyzed effects reflect a spillover from a change in US policies towards China on the performance of Chinese exports to the EU. Moreover, the analysis of the fraction of China's exports shipped to the US supports the assumption of a global fixed cost component.

6 Conclusion

The paper analyzed a potential source of China's export boom to the EU. In contrast to the US, where China benefited from the establishment of permanent normal trade relations, EU

trade policies remained largely unchanged upon China's accession to the WTO. A transmission channel is proposed through which the change in the US trade policy affected China's trade performance at a more general level. 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 the US market will find it easier to export also to the EU. The predictions of the model were confirmed empirically using product-level data at the disaggregated HS6 level. Chinese exports to the EU are about 12 percent higher for goods that were exposed to US tariff uncertainty before WTO entry. In line with the theoretical model, adjustments at the extensive margin could be verified. An analysis of how the effect of the US policy spillover evolves over time suggested that it levels out after a few years.

The findings of this paper bear important implications for the scope of international policy negotiations, as well as for their impact on third countries. To the extent that details about negotiations remain unobserved they may expose firms to unexpected competitive shocks. A closer analysis of the consequences is a possible direction of future research. The paper also presented supportive evidence that exporters face both destination specific and global fixed costs. This complements findings presented by Hanson and Xiang (2011) for services exports, but raises the question as to how this conforms to stylized patterns observed in manufacturing firm-level datasets. Characteristics of the production processes (i.e. standardized vs. differentiated goods) might be one explanation worth pursuing further. A final result revealed in this paper relates to the estimation of model parameters. The estimated effect of the US policy spillover appeared to be correlated with the length of the sample period. A dynamic specification was able to produce more robust quantitative results. With regard to estimating certain parameters of a model, neglection of underlying dynamic processes may lead to serious biases and misinterpretations of the importance of economic variables.

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A   Appendix

A.1   Robustness Checks

This appendix describes data and shows the results of various robustness checks discussed in the main text of this paper.

Unobserved Productivity Effects upon WTO Accession.   The baseline sample used for the analysis of general productivity effects is the same as described in 4.2. The measures for China's RCA were downloaded from the CEPII website (*Centre d'Études Prospectives et d'Informations Internationales*), which provides this information according to the 4-digit HS nomenclature, at the most disaggregated level.[19] Information on exchange rates was drawn from the Penn World Tables 8.0, which provides countries' nominal exchange rates, expressing its domestic currency in US dollars. Exchange cross-rates, i.e. (Yuan/USD)/(ECU/USD), were computed to obtain bilateral exchange rates for China towards European importing countries. An increase of this exchange rate indicates depreciation, and hence easier market access for Chinese exporters. Before re-estimating (14), including the inverted RCA measure, the data was aggregated to the 4-digit HS level. The estimation results shown in Table A1 encompassed 915 4-digit product categories, 14 destinations, and 11 years.

The results shown in Table A2 use information on quality adjusted prices from Feenstra and Romalis (2014).[20] The raw datasets report detailed information exports, imports, unit values, estimated quality, and quality-adjusted prices, among others, for the years 1989-2011, at the 4-digit SITC Rev. 2 level. The data on quality-adjusted prices are expressed relative to a benchmark country, which varies across years, products, and product units. Accordingly, for each product-unit-year the quality-adjusted price for a country was deflated by the corresponding average of all countries. That is, China's relative quality-adusted export price for a SITC-unit pair j, at time t, is $\tilde{p}_{jt}^{cn} = p_{jt}^{cn}/\bar{p}_{jt}$. Relative quality-adjusted import

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[19]http://www.cepii.fr/cepii/en/bdd_modele/presentation.asp?id=26
[20]They can be downloaded at http://cid.econ.ucdavis.edu/Html/Quality_Data_Page.html

Table A1: Chinese Exports to the EU-15 in Comparative Disadvantage Industries; Alternative Samples and Specifications, 1995-2005

Specification Details:	Inv. RCA Interaction			Inv. RCA and Exchange Rate		
	(1)	(2)	(3)	(4)	(5)	(6)
Product Range:	Full	No T&C	Full	Full	No T&C	Full
US Tariff Threat	1.195**	0.779**	1.051**	1.197**	0.782**	1.053**
	(0.151)	(0.177)	(0.151)	(0.150)	(0.177)	(0.151)
EU Tariff	-1.943*	-0.654	-1.338^a	-1.952*	-0.647	-1.346^a
	(0.761)	(0.800)	(0.763)	(0.761)	(0.800)	(0.763)
EU Quota '02			0.069**			0.069**
			(0.006)			(0.006)
EU Quota '05			0.069**			0.069**
			(0.013)			(0.0126)
Inverted RCA	1.199**	1.119**	1.181**	1.164**	1.079**	1.145**
	(0.073)	(0.079)	(0.073)	(0.075)	(0.081)	(0.075)
Inv. RCA×Xrate				0.080**	0.091**	0.0823**
				(0.026)	(0.028)	(0.026)
Observations	90,363	73,534	90,363	90,363	73,534	90,363
R-squared	0.252	0.257	0.254	0.252	0.257	0.255
Fixed effects	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St

Note: Table shows estimates based on Eq. (14), but including also a measure of revealed comparative advantage (RCA) from CEPII, computed for the years 1995-1997. Columns (1)-(3) use the inverted measure interacted with WTO membership, wheras columns (4)-(6) use in addition an interaction of inv. RCA with China's nominal exchange rate to EU-15 member states (see text for further details). Fixed effects: Jn=4-digit product-destination, nt=destination-year, St=sector-year. Robust standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

prices of each EU-15 member state were computed analogously.

In the next step, the unit-specific relative prices were aggregated to the SITC level, using relative export as weights. Changes in relative prices were calculated to identify productivity increases around China's WTO accession. It uses the average relative price of product s at time $t0$ (1997-1999) to deflate its corresponding value at $t1$ (2002-2004). The change in Chinese relative quality adjusted export prices are thus:

$$\Delta \tilde{p}_s^c = \tilde{p}_{st1}^c / \tilde{p}_{st0}^c,$$

which corresponds to the results shown in columns (1) through (3) of Table A2. Columns

(4) through (6) divide this ratio by the corresponding ratio for EU country i's import price:

$$\dot{p}_s^c = \Delta \tilde{p}_s^c / \Delta \tilde{p}_s^i.$$

The price data was matched with the baseline dataset using correspondence tables from the United Nations Statistics Division (UNSD). The estimations were carried out after aggregating the data to the 4-digit SITC level, which resulted in 562 product categories. 2-digit SITC codes reporting any MFA quotas were excluded in columns (2) and (4), respectively. As before, the product-specific price changes were interacted with the WTO indicator variable.

Table A2: Chinese Exports to the EU-15 and Export Price Changes; Alternative Samples and Specifications, 1995-2005

Specification Details:	Price Change Exports			Price Change Exports, deflated		
	(1)	(2)	(3)	(4)	(5)	(6)
Product Range:	Full	No T&C	Full	Full	No T&C	Full
US Tariff Threat	1.243**	0.906**	1.186**	0.974**	0.672**	0.932**
	(0.222)	(0.239)	(0.221)	(0.225)	(0.242)	(0.224)
EU Tariff	-5.878**	-4.163**	-5.536**	-6.037**	-4.415**	-5.771**
	(1.048)	(1.058)	(1.049)	(1.044)	(1.050)	(1.045)
EU Quota '02			0.034**			0.0311**
			(0.006)			(0.006)
EU Quota '05			0.014**			0.013**
			(0.005)			(0.005)
Q-adj. Exp. Price	-1.155**	-1.351**	-1.157**	-0.896**	-1.140**	-0.897**
	(0.131)	(0.143)	(0.131)	(0.105)	(0.119)	(0.105)
Observations	60,115	56,670	60,115	58,963	55,578	58,963
R-squared	0.327	0.323	0.328	0.333	0.328	0.333
Fixed effects	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St

Note: Table shows estimates based on Eq. (14), but includes information on quality adjusted export and import prices from Feenstra and Romalis (2014), between 1997-1999 and 2002-2004. Columns (1)-(3) use changes of relative prices for Chinese exports, wheras columns (4)-(6) deflate those price changes by changes in EU's relative import prices (see text for further details). Fixed effects: Jn=4-digit SITC product-destination, nt=destination-year, St=2-digit SITC sector-year. Robust standard errors in parentheses; significance: [a] $p < 0.1$, [*] $p < 0.05$, [**] $p < 0.01$.

Removal of Investment Restrictions in China. The contract-intensity measures from Nunn (2007) are based on detailed US manufacturing input-output tables in the year 1997.[21] Identifying differentiated and homogeneous input goods, the share of the former in total manufactured input goods produces an industry-specific measure of contract intensity according to the NAICS nomenclature (*North American Industry Classification System*). The identification of differentiated and homogeneous goods is based on the classification of Rauch (1999). Since he groups products into three categories (differentiated, reference priced, and traded on organized exchanges), Nunn (2007) offers two measures of contract intensity. The first uses only the share of differentiated inputs, whereas the second uses the share of differentiated and reference priced goods together.

To match the information on contract intensity with the the HS6 codes of my dataset, I use the correspondece files provided by Pierce and Schott (2009) for the year 1995 (where my dataset starts). NAICS codes in the contract intensity dataset are slightly more aggregated than in the concordance files, so that about 39 percent of the NAICS codes had to be mapped by hand (172 out of 442). Table A3 shows results for the full sample, including both automatically and hand matched products.[22]

Export Subsidies. The inference of Chinese production and export subsidies is indirect. It uses information on products for which the EU initiated anti-dumping (AD) and countervailing duty (CVD) investigations against China.

The data on AD filings is available from Bown (2014), and reports cases initiated by the EU since the 1970s. To select the relevant cases, AD filings against China were selected, if their final decision on domestic industry injury was affirmative, and if the imposed AD measures were not revoked by 1995. The affected products, reported at the 8-digit level, were aggregated to 6-digits and coded equal to one for the years when their investigation started,

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[21]The measures of products contract intensity are available at different levels of aggeregation at http://scholar.harvard.edu/nunn/pages/data-0.

[22]The results for the restricted sample indicate a higher robustness of the US policy spillover, and are available upon request.

Table A3: Chinese Exports to the EU-15 and Products' Contract Intensity; Alternative Samples and Measures, 1995-2005

Contract Intensity Measure:	Differentiated Inputs			Non-homogeneous Inputs		
	(1)	(2)	(3)	(4)	(5)	(6)
Product Range:	Full	No T&C	Full	Full	No T&C	Full
US Tariff Threat	0.593**	0.201*	0.356**	0.571**	0.176^a	0.334**
	(0.076)	(0.090)	(0.077)	(0.076)	(0.091)	(0.077)
EU Tariff	-0.367	0.275	-0.0330	-0.442	0.210	-0.107
	(0.420)	(0.434)	(0.420)	(0.420)	(0.434)	(0.420)
EU Quota '02			0.572**			0.571**
			(0.034)			(0.034)
EU Quota '05			0.436**			0.446**
			(0.049)			(0.049)
Contract Intensity	0.316**	0.286**	0.314**	0.649**	0.546**	0.651**
	(0.076)	(0.083)	(0.076)	(0.115)	(0.124)	(0.115)
Observations	267,631	205,406	267,631	267,631	205,406	267,631
R-squared	0.172	0.177	0.174	0.172	0.177	0.174
Fixed effects	*Jn,nt,St*	*Jn,nt,St*	*Jn,nt,St*	*Jn,nt,St*	*Jn,nt,St*	*Jn,nt,St*

Note: Table shows estimates based on Eq. (14), but including also a measure of products' contract intensity (Nunn, 2007). Columns (1)-(3) use the measure based on differentiated inputs only, wheras columns (4)-(6) use the measure accounting for all non-homogeneous intermediate inputs (see text for further details). Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year. Robust standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.

and unitl the AD measure was revoked. In total, 83 HS6 codes were subject to AD measures in the estimations (Table A4).

The data on CVD is available from Bown (2015), and reports cases initiated by the EU since the 1970s. For China, however, CVD investigations were filed only since 2010. I selected those filings where the final decision on domestic industry injury was affirmative and CVD measures were in force. Although the reporting appeared about one decade after China's WTO entry, their exports may have been subsidized from an earlier time on. To controll for this possibility, the respective HS6 products were interacted with the WTO indicator. In total, HS6 codes were subject to CVD measures in the estimations (Table A4, columns (4)-(6)).

Table A4: Chinese Exports to the EU-15 and Products subject to Anti-dumping and Countervailing Measures; Alternative Samples and Specifications, 1995-2005

Specification Details:	AD Filings			AD Filings & CVD		
	(1)	(2)	(3)	(4)	(5)	(6)
Product Range:	Full	No T&C	Full	Full	No T&C	Full
US Tariff Threat	0.646**	0.233*	0.404**	0.648**	0.233*	0.406**
	(0.076)	(0.091)	(0.077)	(0.076)	(0.091)	(0.077)
EU Tariff	-0.437	0.231	-0.101	-0.437	0.231	-0.102
	(0.421)	(0.435)	(0.420)	(0.421)	(0.435)	(0.421)
EU Quota '02			0.577**			0.576**
			(0.034)			(0.034)
EU Quota '05			0.447**			0.447**
			(0.050)			(0.050)
EU AD Filings	-0.086^a	-0.235**	-0.094*	-0.090^a	-0.234**	-0.097*
	(0.048)	(0.054)	(0.047)	(0.047)	(0.054)	(0.0467)
EU CVD				0.520**	0.310	0.507*
				(0.201)	(0.198)	(0.200)
Observations	268,499	205,966	268,499	268,499	205,966	268,499
R-squared	0.171	0.176	0.173	0.171	0.176	0.173
Fixed effects	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St	Jn,nt,St

Note: Table shows estimates based on Eq. (14), but including also an indicator variable for products subject to anti-dumping (AD) filings and countervailing duties (Bown, 2014, 2015). Columns (1)-(3) use the AD indicator, which equals one for the years in which products were under investigation until the case was closed. Columns (4)-(6) augment the model by interacting products with the WTO indicator, when they became subject to countervailing duties after China's WTO entry (see text for further details). Fixed effects: Jn=product-destination, nt=destination-year, St=sector-year. Robust standard errors in parentheses; significance: ^{a} $p < 0.1$, * $p < 0.05$, ** $p < 0.01$.