The Aftermath of Antidumping: Are Temporary Trade Barriers Really *Temporary*?

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Abstract

This paper provides empirical evidence on the question as to whether antidumping (AD) measures have a negative impact on affected bilateral trade exclusively while being in force, or instead induce lasting negative effects even after their revocation. While we are able to confirm a previously documented trade-depressing effect of AD, our results suggest that affected bilateral trade increases significantly only shortly after the revocation of a previously installed AD measure but there seems to be no further stimulus from the second year on. Consequently, affected trade relations may presumably not fully recover from the trade impediment.

Keywords: Antidumping · temporary trade barriers · trade depression · gravity model of trade · panel data

JEL classification: F13; F14

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1. Introduction

The past 25 years have witnessed ongoing changes in the landscape of global trade policy. While tariffs continue to be the most powerful trade barrier in force, their average has fallen within this period both multilaterally in the course of the World Trade Organization’s (WTO) Most-Favored-Nation practice, and more regionally owing to the contemporary growth of regional trade agreements (RTAs). Compared to this, there has been a considerable increase in the use of non-tariff barriers with antidumping (AD) activity being the most prominent one (Prusa, 2005; Bown, 2011; Kee *et al*., 2013).

AD policy has long been the playground of only a handful of mainly developed countries led by the United States, the European Union, and Canada. Part of the extensive increase in global AD activity beginning from the 1990s is attributable to so-called “new heavy users”, a group of emerging market economies (Bown, 2008). To this date, however, the stimulus for this proliferation remains controversial. While AD policy finds WTO-legitimation if applied in order to prevent or offset “unfair” price setting in international trade relations, empirical literature has examined and partly found supportive evidence for various alternative —rather protectionist-inspired— explanatory approaches ranging from “trade policy substitution-” (Feinberg and Reynolds, 2007; Bown and Tovar, 2011; Moore and Zanardi, 2011; Ketterer, 2016) to “retaliation-” (Blonigen and Bown, 2003; Feinberg and Reynolds, 2006; Vandenbussche and Zanardi, 2008) to “safety valve” (Niels and ten Kate, 2006; Moore and Zanardi, 2009) motives.

Not least the expanded circle of applicants has generated a large body of empirical literature regarding the *immediate* impact of AD measures on affected bilateral trade that consistently detects a negative relation in this respect.[1] While the majority of papers has analyzed

[1] Blonigen and Prusa (2016) give a comprehensive literature review of the trade effects associated with AD policy.

affected trade relations at sectoral levels (Staiger and Wolak, 1994; Prusa, 1997; Brenton, 2001; Konings et *al.*, 2001; Prusa, 2001; Durling and Prusa, 2006; Ganguli, 2008; Park, 2009), there is also evidence that AD activity hampers aggregate bilateral trade (Vandenbussche and Zanardi, 2010; Egger and Nelson, 2011).

Albeit of equally high relevance from a trade policy perspective, a topic that has received only marginal academic attention concerns the question of what happens to bilateral trade relations after the removal of previously imposed AD protection. On these grounds, in this paper we attempt to assess as to whether AD measures —that are classified as temporary trade barriers— have a previously documented negative impact on affected bilateral trade (trade depression) exclusively while being in force, i.e. in merely a *temporary* manner, or instead induce *lasting* negative implications for affected bilateral trade even after their revocation. Insights from our research may be considered important particularly with regard to the appropriately targeted implementation of trade policy that is of relevance not only with respect to AD policy, but also for the implementation of other non-tariff barriers or even tariff policy.

Our research intention is exclusively an empirical one. Nevertheless, theoretical intuition may offer explanations for both abovementioned outcomes. On the one hand, the removal of AD protection presumably (re-)stimulates bilateral trade just as straightforward as the well-studied elimination or reduction of tariffs owing to an immediate improvement in market access conditions of previously targeted exporters. On the other hand, however, the burden of a temporary trade barrier may prompt targeted countries to shift their export activities to other destinations that offer comparatively freer market access in the meantime. If entering into new markets required fixed market entry costs the reorientation of exports could persist, and with this we may find lasting negative implications for trade relations between imposing and target countries of AD protection.[2] Moreover, even if an AD measure is finally removed, the fact that an AD measure had been installed at some point in time might have a signaling effect concerning the willingness of a country to impose a renewed AD measure. This poses a risk for the previously targeted country and might in turn lead to distorted price setting and export quantities which both potentially impede trade (Pauwels *et al.*, 2001). As Egger and Nelson (2011) argue, there is considerable anecdotal but not systematic evidence that previously targeted countries adjust their pricing in order to avoid the risk of AD.

To the best of our knowledge, Bown (2013) is the only previous publication that empirically addresses the impact of the removal of temporary trade barriers on bilateral trade. Analyzing a total of 746 AD cases that have been imposed and removed across fourteen G20 member countries between 1992 and 2008 in a linear regression framework, the author discovers that exports (volumes and shares alike) of targeted countries experience an immediate increase once being relieved from the trade barrier. In particular China's export performance is found to adjust comparatively quickly and aggressively to improving market conditions. With the exception of China, however, targeted countries generally could not manage to re-establish pre-AD export levels. One may interpret this finding as a first indication for a hysteresis effect of AD activity.

Our paper deviates from Bown (2013) in a number of aspects. Firstly, we extend both the period of investigation (from 1991 to 2014), and the number of imposing and targeted countries (31 countries, including both developed and developing countries) which allows to consider roughly two and a half times more AD cases. Secondly, following large parts of previous literature on the immediate trade impact of AD activity, we employ a gravity-inspired dynamic panel data model that allows to control for forms of unobserved heterogeneity, and in which we also take into account two variables that have been attested to significantly im-

[2] Albeit both a direct theoretical link as well as empirical evidence are lacking for the latter line of argumentation with respect to AD activity to date, previous work on a permanent or hysteresis impact of exchange rate shocks on bilateral trade patterns may well provide support in this respect.

pact bilateral trade, namely tariff policy and exchange rate volatility. In this respect, the results obtained by Bown (2013) might suffer from a serious bias. Lastly, considering that the potential existence of lasting trade depression may well depend on the intensity and dimension of the initial trade policy shock we explicitly analyze the aftermath of AD activity on bilateral trade with respect to different case durations. Our findings indicate that AD policy may have a more concerning trade-depressing effect than previous studies have suggested, especially for lower income countries.

The remainder of this paper is structured as follows: Along with a detailed introduction of the data, our methodology is outlined in section 2. Section 3 presents the main results whereas section 4 discusses various extensions of our basic model. Section 5 concludes with policy implications.

2. Data and Methodology

Data

We extract annual bilateral import flows at the Harmonized System (HS) four-digit commodity level from the UN Comtrade database and combine those with aggregated AD information from the World Bank's Global Antidumping Database (GAD) collected by Bown (2015) for a symmetric bilateral panel of 31 countries (see Appendix 1 for the list of countries in our sample). The overall time period under consideration ranges from 1991 to 2014. For several imposing countries, however, AD data are available only beginning from later years than 1991. Compared to this, our data source also contains information regarding pre-1991 AD activity for individual countries, allowing to partly retrace four-digit AD cases that have come into force prior to 1991. With this, we are able to record the precise total duration of four-digit AD cases in our sample whenever available.

We take into account only those country-pair-commodity combinations that have experienced bilateral AD activity in the years between 1991 and 2014 with a minimum duration of at least twelve months, i.e. a full calendar year.[3] As we consider annual instead of monthly data for our empirical analysis this translates into a minimum duration of at least three consecutive dataset-years since an AD case may begin/end at any point in time throughout the calendar year.

Additionally, we collect AD cases independent of being declared as a preliminary or a final measure. This *in*distinction brings both disadvantages and advantages. While on the one hand, the generalization does not allow to analyze separately the effects of preliminary and final AD measures on bilateral trade relations. On the other hand, exclusive concentration on final AD measures would distort precise identification of resulting trade effects due to the omission of preceding trade barriers in force. In this respect, Staiger and Wolak (1994) give evidence on a nearly identical impact of preliminary and final AD measures on trade flows.

At a more detailed inspection, our sample originally covers a total of 3,058 bilateral four-digit AD cases between 1991 and 2014, i.e. (i) those that have been initiated prior to 1991 and have still been in force beyond 2014, (ii) those that have been initiated prior to 1991 and have been revoked by 2014, (iii) those that have been initiated in or after 1991 and have still been in force beyond 2014, and lastly (iv) those that have been initiated in or after 1991 and have been revoked by 2014. As we attempt to shed light on the question of what happens to bilateral trade relations once relieved from previously imposed AD protection, out of the total of 3,058 cases, we consider only those for empirical analysis that have not been in force beyond 2014 —in other words, only those that have been terminated between 1991 and 2014— so that our sample reduces to 1,844 bilateral four-digit AD cases. Their mean duration is 9.01 years. Table 1 gives an indication of the frequency of AD cases by their duration in years.

[3] We consider that AD activity that endures less than twelve months does not point towards a sustainably pursued trade policy. As regards our sample there are, however, only a handful of AD cases that had a shorter duration than twelve months.

{Please insert Table 1 about here}

As can be seen, case durations are far from being normally distributed where a cumulated frequency of 50 percent is found for those with a maximum in-force duration of seven years. When examining the noticeable peak of AD cases with a duration of 16 years, Mexico alone accounts for 192 out of the 216. Remarkably, all respective Mexican AD measures were installed from 1993 onwards.[4] As concerns the eight outliers with a duration of 24 years each, all of them refer to the United States as imposing country. Targeting Japan and the European Union, seven cases span a period from 1988 to 2011 whereas one has been initiated in 1987 and revoked in 2010.

Moreover, Table 2 lists the ten most frequent AD users within our sample that account for 86.1 percent of respective total AD cases. The group is composed of so-called "traditional-" (United States, European Union, Canada, and Australia) and a number of "new heavy" (Mexico, India, Argentina, South Africa, Brazil, and Peru) AD users. Among them, Mexico stands out not only with respect to the number of cases but also with an average imposition of case-specific AD measures of 13.9 years.

{Please insert Table 2 about here}

Our dataset includes annual effectively applied trade-weighted bilateral tariff rates at the HS four-digit commodity level taken from the World Integrated Trade Solutions (WITS) database whenever readily available.

Lastly, we contribute to the existing literature by incorporating the real exchange rate that we calculated using UNCTAD data on annual nominal exchange rates (NER)[5] and Consumer Price Index (CPI) data obtained from UNCTAD statistics (primary source) and the World

[4] This coincides with previous evidence on rocketing Mexican AD activity in 1993 such as in Francois and Niels (2003) and Niels and ten Kate (2004) where the former gives Mexico's appreciations of its real exchange rate as a possible explanation.

[5] Nominal exchange rates towards US-Dollar are used to derive bilateral exchange rates for each country with regard to any other country.

Bank.[6] Taking into account the European integration process, we calculated a representative exchange rate for the European Union with respect to its trading partners. The methodology and computational steps are provided in Appendix 4.

Estimation Issues

For econometric implementation, we employ a simple gravity-inspired dynamic panel data model as our baseline specification that reads as follows:

$$\begin{aligned} \ln(M)_{ijkt} = \alpha_0 + \beta_1 \ln(M)_{ijkt-1} + \beta_2 \ln(M)_{ijkt-2} + \beta_3 \ln\left(1 + t_{ijkt}\right) + \beta_4 \ln\left(RER_{ijt}\right) \\ + \sum_{n=T-3}^{T-1} \gamma_n \,\text{before}_{ijk,n} \\ + \sum_{m=T+1}^{T+3} \delta_m \,\text{after}_{ijk,m} + \pi_{ijk} + \eta_{it} + \mu_{jt} + \varphi_{kt} + u_{ijkt} \end{aligned} \quad (1)$$

where M_{ijkt} denotes country i's imports from j in HS four-digit sector k and period t, α_0 is a constant, and u_{ijkt} is the error term that is assumed to be well behaved with $\sim iid\ (0, \sigma_u^2)$. In order to capture bilateral import dynamics, we include the first as well as the second lag of the dependent variable. Moreover, t_{ijkt} is a continuous variable indicating the absolute ad-valorem effectively applied tariff of i towards j in commodity k and period t. RER_{ijt} is the real exchange rate between i and j in period t and a rise equals a depreciation of the importer's currency which translates into a loss of purchasing power abroad, resulting in a reduction of imports.

In addition, we incorporate country-pair-commodity effects (π_{ijk}) in order to control for unobserved heterogeneity among triads, and importer-year- and exporter-year effects (η_{it} and μ_{it}, respectively) to account for time-varying omitted variables such as multilateral trade resistance following Baier and Bergstrand (2007). The inclusion of country-pair-commodity effects only allows an interpretation of the within-group effects. As there may be annual

[6] CPI data is not available for China for the years 1991 and 1992. For this reason, the trend in the GDP deflator (annual percent changes) was used to fill the data series.

commodity-specific global trends in bilateral trade, we also control for time-varying commodity effects through φ_{kt}.

While the inclusion of π_{ijk} resolves a potential bias stemming from the correlation of regressors and unobserved commodity-country-pair specific heterogeneity, there remains a correlation between the lagged dependent variables and the error term because the former are a function of the latter (Nickell, 1981). However, Roodman (2009) argues that the bias in question will become insignificant the larger the T-dimension in panels. For the dataset at hand, which is covering a time span of 25 years, *classical* Ordinary Least Squares (OLS) fixed effects thus appears to be comparatively more efficient than the System Generalized Method of Moments (GMM) estimator suggested by Arellano and Bover (1995) and Blundell and Bond (1998).[7]

Our research interest concentrates on the trade-impact of an AD measure's transition from its status of being "in force" to being "revoked". Two sets of policy dummy variables therefore form the heart of equation (1) where we define period T as the revocation year of an AD measure imposed by country i towards j in sector k. Then, the dummy variable set $\text{before}_{ijk,n}$ creates three different policy dummy variables that equal unity for the last in-force year of an AD measure ($\text{n} = T - 1$), and for its previous two years ($\text{n} = T - 2$, and $\text{n} = T - 3$, respectively), all zero otherwise.[8] Importantly, depending on the actual duration of individual AD cases, $\text{before}_{ijk,n}$ may not only refer to the *last* years of an AD measure in force but could likewise capture its *initial* years. In a similar vein, $\text{before}_{ijk,n}$ reduces to a two-period specification for those AD cases with a maximum duration of three years. Recapitulating empirical findings regarding the trade-depressing impact of AD activity on affected bilateral trade, nonetheless, we suspect coefficient estimates of all three to be negatively signed.

[7] However, the GMM coefficient estimates of the before and after dummy variables confirm the results obtained from the FE regression model presented below. Results are available on request.
[8] Since our main interest is to investigate what happens when an AD measure is revoked and not when it is put in place, we refrain from using an implementation dummy as done in Bown (2013).

By comparison, the dummy variable set $\text{after}_{ijk,m}$ creates three different policy dummy variables that equal unity for the first full year after the revocation of a previously imposed AD measure ($\text{m} = T + 1$), and for the following two years ($\text{m} = T + 2$, and $\text{m} = T + 3$, respectively), all zero otherwise. On the one hand, its coefficient estimates may be negatively signed, signaling lasting trade depression even after an AD measure's revocation. On the other hand, its coefficient estimated could likewise be positively signed, indicating a boost in bilateral trade once the trade barrier has been removed.

An alternative empirical strategy to analyze the temporal effects of AD on bilateral trade compared to employing dummy variables would be to consider AD activity more *explicitly*, i.e. taking into account its actual degree of duty. However, there are several caveats speaking against it: (i) Firstly, the type of the AD measure installed may vary widely across cases encompassing ad-valorem duties, specific duties, suspension agreement, price undertaking from exporters, or duties becoming due only if export prices fall under a certain threshold, which makes consistent incorporation difficult; (ii) Secondly, albeit we use the most comprehensive database currently available for obtaining bilateral AD information, there are nonetheless numerous missing or inconsistent data entries when it comes to duty amounts; (iii) Thirdly, as our main interest lies upon the transition period of bilateral AD measures from being in force to being revoked, one would find strictly *zero* duties once a measure had been revoked which implies no further informative benefit when incorporated continuously instead of binary; (iv) Lastly, as discussed in Prusa (2001), targeted countries may respond to imposed AD duties by adjusting their export prices so that a continuous variable capturing AD activity would presumably underestimate its actual impact on affected bilateral trade.

3. Main Results

Table 3 presents the results of our baseline specification. Columns (1) and (2) report coefficient estimates for model specifications that separately include the sets of the before- and after policy dummy variables, respectively, and column (3) displays those for their joint incorporation. Across model specifications, both the first and the second lag of the dependent variable are statistically significant at the 0.1 percent level. Their coefficient estimates are positively signed as is expected and close to those obtained in previous works (compare e.g. Ganguli, 2008 and Park, 2009). While sign and magnitude of the effectively applied bilateral tariff and the real exchange rate are in line with expectations and the gravity literature, however, the estimates are statistically insignificant possibly due to the incorporation of importer-year-, exporter-year- and commodity-year effects that already absorb a lot of temporal variation across countries and sectors.

{Please insert Table 3 about here}

Turning towards AD activity, we focus on the results obtained from the joint estimation reported in column (3) and only briefly compare coefficient estimates to those of the restricted model versions reported in columns (1) and (2).

As expected, the policy dummy variables capturing the last three in-force years of an AD measure prior to its revocation are negatively signed and significant (at least at the five percent level). We observe the largest effect in the last year before the trade barrier is removed ($T-1$) where in economic terms, based on its coefficient estimate, affected bilateral imports are reduced by some eleven percent in the last in-force year of an AD measure compared to a hypothetical counterfactual.[9] Although smaller in economic terms and with a lower level of statistical significance, affected bilateral imports also appear to be lower in $T-2$ and $T-3$ as a result of the trade barrier (6.5 percent and 7.9 percent, respectively). Note that in the dataset at hand, $\text{before}_{ijk,T-3}$ may also capture the initial year of the trade barrier what may result in a higher coefficient estimate in $T-3$ than in $T-2$.

The comparatively stronger negative effect in terms of magnitude in the last in-force year of AD could find an explanation in a possible anticipation effect of market participants in imposing countries. Although there is no general rule for public notification when it comes to the revocation of a previously installed AD measure, market participants in imposing countries may well have information on an imminent revocation, exemplarily owing to a reopening of the AD investigation in question or even official announcements in this respect. In expectation of a reduction of import prices as a result of future relieves from the AD burden, domestic consumers could as a consequence be motivated to postpone their consumption of foreign goods and show an even lower import demand in the meanwhile.

When inspecting coefficient estimates of the after dummy variables, we find that affected bilateral imports increase in the very first year after the removal of the previously installed AD measure ($T+1$). The effect is estimated to be statistically significant at the five percent level. The coefficient estimate indicates that imports are 6.5 percent higher compared to the counterfactual of an ongoing AD case.[10] However, in $T+2$ and $T+3$ previously affected bilateral imports do not show any significant positive reaction to the trade barrier's removal anymore. Likewise, we neither find a respective significant effect for the (unreported) fourth year after revocation. Moreover, we also included a policy dummy variable that equals unity

[9] Marginal effects of dummy variables' coefficient estimates are calculated as $\left(\exp\left\{\hat{\beta}_k - \frac{1}{2}\hat{V}\left(\hat{\beta}_k\right)\right\} - 1\right) \times 100$ following Kennedy (1981), where $\hat{\beta}_k$ is a dummy variable's coefficient estimate and $\hat{V}\left(\hat{\beta}_k\right)$ its variance.

[10] The dummy variable equals zero likewise if the AD measure has not been in force yet. However, note that the inclusion of country-pair-commodity fixed effects only allows an interpretation of the within-group effects, hence a change over time. Therefore, the after dummy variable is mainly interpreted as a change from zero (AD still in force) to one (AD revoked).

in the year of revocation (period *T*) but did not obtain any significant effect in doing so. The findings for the after dummy variables point at a relatively weak trade-resuming effect in contrast to a strongly negative effect of a previously installed measure.

We obtain only slightly different results when the before- and after policy dummy variables are separately included (columns (1) and (2)). All coefficient estimates are slightly larger than those reported in column (3) und the after dummy variable in $T + 2$ is significant at the ten percent level. However, also regarding the higher significance level of the before policy dummy variable in $T - 2$, findings of our baseline results are confirmed throughout.

To put in a nutshell, we find statistically significant negatively signed coefficient estimates for *all* before dummy variables but only *one* positively signed and statistically significant coefficient estimates for the after dummy variables. With this, we cautiously conclude that the reduction of affected imports due to the imposition of an AD measure can presumably not be compensated entirely and that lasting trade-depressing effects arising from the trade barrier are very much likely.

4. Extension and Robustness Checks

Differentiation by Case Durations

According to WTO law (Article VI, Art. 11) and reflecting the literal meaning of a temporary trade barrier, *de jure* AD measures should be terminated not later than five years after their imposition. *De facto*, this may nevertheless not necessarily be the case as indicated by the above shown variation in case durations in Table 1. Related to this one might assume that the effect on trade flows differs with respect to the duration of the AD measure. Trade relations affected throughout a comparatively longer time span will presumably only be revived after a certain time – if at all. Targeted countries may have shifted their export activities to other destinations and fostered those new trade relations in the meanwhile, and hence it takes more time (here three years) for previously affected bilateral imports to pick up again.

Therefore, we speculate that longer case durations will lead to a higher likelihood of lasting trade depression in affected bilateral trade and separately analyze short-run (shorter than six years) and long-run (longer than twelve years) cases. Avoiding a sample split, we generate four separate sets for each of the before- and after policy dummy variables in the same manner as in the baseline regression yet involving a further differentiation with respect to those case durations (i) shorter than six years, (ii) longer or equal than six years, (iii) shorter than twelve years, and (iv) longer or equal than twelve years. Results are presented in Table 4.

For those AD cases that are in force up to five years (column (1)) we cannot find any effect on bilateral trade in the years before and after revocation. With this, the baseline regression results presented in Table 3 seem to be driven primarily by cases that are in force longer or equal than six years. Hence, AD measures that are literally in force only "temporarily" do not seem to affect bilateral imports significantly.[11]

{Please insert Table 4 about here}

The results reported in columns (2) and (3) where regressions include cases that are in force longer or equal than six and shorter than twelve years, respectively, are very much in line with the baseline results. We again find affected bilateral imports to be significantly reduced in the last in-force years of an AD measure. A similar picture emerges when we turn towards the trade effects associated with the removal of a previously installed AD measure. As is the case with the immediate impact of AD on affected bilateral trade, we find a continuous trade-depressing effect of AD on affected bilateral trade and a statistically significant stimulation only shortly ($T + 1$) after the revocation.

[11] In this setting, the $\text{before}_{ijk,T-3}$ dummy variable certainly captures the initial year of the three-year cases in our sample but is also not statistically significant.

As concerns those cases with a duration of twelve or more years, we only observe a small trade-depressing effect two years before the AD is finally revoked. In addition, trade does not seem to resume immediately after the revocation but instead rather slightly after three years.

In summary, the baseline regression results seem to be driven by cases that are in force longer or equal than six but also shorter than twelve years.

Differentiation by Income Levels of Imposing Country

We presume that the imposition of AD may be more trade depressing (in both its immediate and its lasting impact) for countries that provide only marginal economic attraction for exporting countries as is the case with most developing countries. In this regard, one may assume that targeted exporting countries could maintain a reorientation of their export activities even after the revocation of the AD measure in question due to reasons of little market potential of previous trading partners.

Therefore, we estimate the baseline model specification separately for high-income countries (Table 5 columns (1) through (3)) and middle and low-income countries (Table 5 columns (4) through (6))[12]. In the following discussion of the results we again focus on our main specifications reported in column (3) and (6) in Table 5. The results for high income countries (19,087 observations) are similar to the baseline regression results as concerning the magnitudes of coefficient estimates and findings for the after dummy variables, with the exception that we find the before variables to be significant only in *T-1*. In contrast, an AD measure imposed by lower income countries has a significantly negative effect on affected bilateral imports across all three years before the AD is revoked but no significant trade-resuming effect. We cautiously argue that trade barriers harm high-income countries' imports less than lower income countries' imports.

[12] The classification of high, middle and low-income countries is based on the World Bank Atlas methodology. In 2003, the threshold was 765 US-Dollars for low income, 3,035 US-Dollars for lower middle income, 9,385 US-Dollars for upper middle income and above 9,385 US-Dollars high income countries.

{Please insert Table 5 about here}

Share of Imports as Dependent Variable

Lastly, we re-estimate equation (1) with country *j*'s share in *i*'s total sector *k* imports in period *t* (ln M_SHARE_{ijkt}) as the dependent variable instead of the logarithm of bilateral imports. Shares are frequently used as a dependent variable in the literature to examine trade diversion effects arising from AD activity (Park, 2009; Brenton, 2001). Within our sample, a lower country *j*'s share in total *i*'s import implies that, *ceteris paribus*, imports are diverted to other markets. Hence, we assess the effect of an AD measure on changes in the economic importance of one country with respect to total imports. We find that an AD measure in its last three years of implementation significantly reduces the importance of the targeted country. Although the coefficients are slightly smaller, the results indicate that a loss of imports by 10.1 percent (Table 6, column (3)) in the last year of in force cannot be compensated by an increase of 6.4 percent attributed to the revocation. Most importantly, we detect three significantly negatively signed before policy variables but only one positively signed after variable. In summary, we quantitatively confirm that the burden of a temporary trade barrier may indeed prompt targeted countries to shift exports to other destinations in the meantime. We find lasting negative implications for trade relations between imposing and target countries of AD protection.

{Please insert Table 6 about here}

5. Concluding Remarks

In the past decades, AD policy has evolved into a widely applied trade barrier of both developed and developing countries. While its original intention lies in the prevention or offset of price dumping in international trade, recent economic literature has given rise to the concern that AD may simply be another protectionist measure in the set of trade policies. In this paper, we have not only examined the question as to whether AD measures have the previously documented negative impact on affected bilateral trade exclusively while being in force but also whether they have lasting negative implications even after their revocation.

Building upon a gravity-inspired dynamic panel data model, our results reveal several insights. Firstly, we find a continuously strong trade-depressing effect of AD on affected bilateral trade even shortly before the revocation of the trade barrier. Secondly, affected bilateral trade increases significantly in the first full year after the revocation of a previously installed AD measure. However, we do not find evidence of a respective stimulus from the second year following the revocation. The estimated effects seem to be primarily driven by those AD cases that have a duration of more than six but less than twelve years. Differentiating the sample into high-income and low- and middle-income countries indicates that the trade-depressing effect is much more pronounced in low- and middle-income countries whereas the increase of trade after the revocation of the AD measure can only be observed for high-income countries. The results also hold for the use of share of imports as dependent variable instead of the import value. Additionally, they are robust to the inclusion of panel fixed- and various individual effects.

Based on our findings, we argue that presumably the reduction of imports that a country experiences in the course of the imposition of an AD measure is not compensated by its revocation. Conversely, the negative effect of AD measures on affected bilateral trade is not restricted to their actual temporal duration. Instead, lasting negative implications arising from

AD policy on affected bilateral trade are highly probable. One possible reason for these findings could be that in response to AD measures targeted countries permanently shift some of their export activities to other countries. The observation that the effect is most pronounced for case durations longer than five years seems plausible since a longer duration increases the incentive for countries to find alternative trade partners. Moreover, low- and middle-income countries seem to face stronger lasting trade-depressing effects. This could be due to their limited economic importance and limited market potential and/or to the higher elasticity of substitution and resulting higher competitive pressure.

From a global trade policy perspective AD measures may thus be considered even more concerning in terms of a trade-depressing effect than previous studies have suggested, especially for lower income countries. With this, it should be chosen as a policy option more cautiously than is currently the case.

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Table 1. Frequency of Bilateral Four-digit Antidumping Cases by Duration (Terminated Cases Only), 1991-2014

Duration in Years	*Number of Antidumping Cases*
3	80
4	107
5	82
6	485
7	278
8	98
9	42
10	55
11	106
12	92
13	60
14	31
15	38
16	216
17	20
18	4
19	14
20	9
21	11
22	8
24	8
Total	1,844

Notes: Duration in years refers to dataset-years. AD data based on countries considered for empirical analysis (see section 3 for details). Own calculation based on data from Bown (2015).

Table 2. Most Frequent Four-digit Antidumping Users (Terminated Cases Only), 1991-2014

Imposing Country	*Number of Antidumping Cases*	*Average Duration in Years*
Mexico	272	13.9
United States	259	11.3
European Union	229	8.0
India	198	6.7
Canada	172	9.6
Argentina	141	6.0
Australia	115	7.1
South Africa	86	7.9
Brazil	71	7.6
Peru	44	8.5
Total	1,587	9.32

Notes: Duration in years refers to dataset-years. AD data based on countries considered for empirical analysis (see section 3 for details). Own calculation based on data from Bown (2015).

Table 3. Baseline Model Specification — Panel Fixed Effects

VARIABLES	(1)	(2)	(3)
$\ln(M)_{ijkt-1}$	0.561***	0.561***	0.561***
	(0.0164)	(0.0164)	(0.0164)
$\ln(M)_{ijkt-2}$	0.0646***	0.0653***	0.0650***
	(0.0130)	(0.0130)	(0.0130)
$\ln(1 + t_{ijkt})$	-0.162	-0.194	-0.172
	(0.266)	(0.269)	(0.267)
$\ln(RER_{ijt})$	-0.0497	-0.0459	-0.0517
	(0.0913)	(0.0913)	(0.0915)
$\text{before}_{ijk,T-1}$	-0.124***	—	-0.116***
	(0.0271)		(0.0274)
$\text{before}_{ijk,T-2}$	-0.0745**	—	-0.0674*
	(0.0269)		(0.0272)
$\text{before}_{ijk,T-3}$	-0.0896**	—	-0.0821**
	(0.0294)		(0.0294)
$\text{after}_{ijk,T+1}$	—	0.0827**	0.0634*
		(0.0272)	(0.0276)
$\text{after}_{ijk,T+2}$	—	0.0450†	0.0261
		(0.0254)	(0.0256)
$\text{after}_{ijk,T+3}$	—	0.0243	0.00705
		(0.0260)	(0.0260)
Observations	36,880	36,880	36,880
RMSE	0.724	0.725	0.724
R^2 adj. within	0.384	0.384	0.384
R^2 adj. overall	0.925	0.925	0.925

Notes: Robust, clustered (at the country-pair-commodity level) standard errors in parentheses. Country-pair-commodity- (π_{ijk}), importer -year- and exporter- year effects, and commodity-year effects (η_{it}, μ_{jt} and φ_{kt} respectively) are always included in the estimation but not reported. Asterisks denote the level of statistical significance with *** p<0.001, ** p<0.01, * p<0.05, † p<0.1.

Table 4. Model specification by case duration (cd) — Panel Fixed Effects

	(1)	(2)	(3)	(4)
VARIABLES	cd < 6 years	cd ≥ 6 years	cd < 12 years	cd ≥ 12 years
$\ln(M)_{ijkt-1}$	0.561***	0.561***	0.561***	0.561***
	(0.0164)	(0.0164)	(0.0164)	(0.0164)
$\ln(M)_{ijkt-2}$	0.0647***	0.0650***	0.0650***	0.0647***
	(0.0130)	(0.0130)	(0.0130)	(0.0130)
$\ln(1+t_{ijkt})$	-0.185	-0.177	-0.192	-0.163
	(0.267)	(0.267)	(0.266)	(0.269)
$\ln(RER_{ijt})$	-0.0414	-0.0539	-0.0470	-0.0453
	(0.0913)	(0.0910)	(0.0915)	(0.0909)
$\text{before}_{ijk,T-1}$	-0.0203	-0.125***	-0.137***	-0.0636
	(0.0769)	(0.0288)	(0.0328)	(0.0535)
$\text{before}_{ijk,T-2}$	-0.0497	-0.0688*	-0.0643*	-0.101†
	(0.108)	(0.0277)	(0.0314)	(0.0520)
$\text{before}_{ijk,T-3}$	-0.0587	-0.0839**	-0.0803*	-0.0409
	(0.120)	(0.0298)	(0.0364)	(0.0497)
$\text{after}_{ijk,T+1}$	0.0538	0.0649*	0.0726*	0.0266
	(0.0728)	(0.0295)	(0.0308)	(0.0583)
$\text{after}_{ijk,T+2}$	0.00587	0.0295	0.0213	0.0497
	(0.0550)	(0.0287)	(0.0290)	(0.0559)
$\text{after}_{ijk,T+3}$	0.0732	-0.00371	-0.00941	0.0975†
	(0.0625)	(0.0283)	(0.0283)	(0.0588)
Observations	36,880	36,880	36,880	36,880
RMSE	0.725	0.724	0.724	0.725
R^2 adj. within	0.383	0.384	0.384	0.383
R^2 adj. overall	0.925	0.925	0.925	0.925

Notes: Robust, clustered (at the country-pair-commodity level) standard errors in parentheses. Country-pair-commodity- (π_{ijk}), importer -year- and exporter- year effects, and commodity-year effects (η_{it}, μ_{jt} and φ_{kt} respectively) are always included in the estimation but not reported. Asterisks denote the level of statistical significance with *** p<0.001, ** p<0.01, * p<0.05, † p<0.1.

Table 5: Results by income group

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	High-income countries	High-income countries	High-income countries	Low- and middle-income countries	Low- and middle-income countries	Low- and middle-income countries
$\ln(M)_{ijkt-1}$	0.561***	0.561***	0.561***	0.541***	0.542***	0.541***
	(0.0272)	(0.0273)	(0.0272)	(0.0218)	(0.0218)	(0.0218)
$\ln(M)_{ijkt-2}$	0.0872***	0.0886***	0.0877***	0.0628***	0.0635***	0.0636***
	(0.0207)	(0.0208)	(0.0208)	(0.0175)	(0.0175)	(0.0175)
$\ln(1+t_{ijkt})$	-0.488	-0.513	-0.506	-0.302	-0.349	-0.308
	(0.558)	(0.556)	(0.556)	(0.438)	(0.440)	(0.438)
$\ln(RER_{ijt})$	-0.0587	-0.0699	-0.0631	-0.0592	-0.0417	-0.0610
	(0.148)	(0.149)	(0.148)	(0.144)	(0.143)	(0.144)
$before_{ijk,T-1}$	-0.160***	—	-0.153***	-0.121**	—	-0.107*
	(0.0422)		(0.0426)	(0.0433)		(0.0448)
$before_{ijk,T-2}$	-0.0324	—	-0.0277	-0.146**	—	-0.133**
	(0.0384)		(0.0385)	(0.0470)		(0.0478)
$before_{ijk,T-3}$	-0.0481	—	-0.0430	-0.189***	—	-0.175**
	(0.0376)		(0.0382)	(0.0539)		(0.0539)
$after_{ijk,T+1}$	—	0.0933**	0.0773*	—	0.0995†	0.0648
		(0.0348)	(0.0352)		(0.0514)	(0.0529)
$after_{ijk,T+2}$	—	0.0308	0.0177	—	0.106*	0.0700
		(0.0379)	(0.0384)		(0.0445)	(0.0449)
$after_{ijk,T+3}$	—	-0.0103	-0.0228	—	0.0693	0.0397
		(0.0366)	(0.0365)		(0.0494)	(0.0497)
Observations	19,087	19,087	19,087	15,456	15,456	15,456
RMSE	0.682	0.682	0.682	0.773	0.773	0.773
R^2 adj. within	0.404	0.404	0.404	0.350	0.349	0.350
R^2 adj overall	0.929	0.929	0.929	0.904	0.903	0.904

Notes: Robust, clustered standard errors in parentheses. Country-pair-commodity- (π_{ijk}), importer -year- and exporter- year effects, and commodity-year effects (η_{it}, μ_{jt} and φ_{kt} respectively) are always included in the estimation but not reported. Asterisks denote the level of statistical significance with *** p<0.001, ** p<0.01, * p<0.05, † p<0.1.

High-income countries: Australia, Canada, EU, Israel, Korea, New Zealand, Taiwan, USA

Low- and middle-income countries: Argentina, Brazil, Chile, China, Colombia, Costa Rica, India, Indonesia, Mexico, Malaysia, Pakistan, Peru, Philippines, South Africa, Thailand, Trinidad and Tobago, Turkey, Venezuela

Table 6: Model Variation — Import share as dependent variable

VARIABLES	(1)	(2)	(3)
ln M_SHARE_{ijkt-1}	0.547***	0.548***	0.547***
	(0.0163)	(0.0163)	(0.0163)
ln M_SHARE_{ijkt-2}	0.0720***	0.0726***	0.0724***
	(0.0131)	(0.0131)	(0.0131)
$\ln(1+\mathrm{t_{ijkt}})$	-0.101	-0.129	-0.110
	(0.241)	(0.242)	(0.241)
$\ln(RER_{ijt})$	-0.0419	-0.0390	-0.0438
	(0.0872)	(0.0873)	(0.0874)
$\mathrm{before_{ijk,T-1}}$	-0.114***	—	-0.107***
	(0.0262)		(0.0264)
$\mathrm{before_{ijk,T-2}}$	-0.0676**	—	-0.0607*
	(0.0256)		(0.0260)
$\mathrm{before_{ijk,T-3}}$	-0.0700*	—	-0.0628*
	(0.0283)		(0.0283)
$\mathrm{after_{ijk,T+1}}$	—	0.0797**	0.0628*
		(0.0262)	(0.0266)
$\mathrm{after_{ijk,T+2}}$	—	0.0384	0.0220
		(0.0246)	(0.0248)
$\mathrm{after_{ijk,T+3}}$	—	0.0230	0.00795
		(0.0252)	(0.0252)
Observations	36,880	36,880	36,880
RMSE	0.705	0.705	0.705
R^2 adj. within	0.373	0.372	0.373
R^2 adj. overall	0.872	0.872	0.872

Notes: Robust, clustered standard errors in parentheses. Country-pair-commodity- (π_{ijk}), importer -year- and exporter- year effects, and commodity-year effects (η_{it}, μ_{jt} and φ_{kt} respectively) are always included in the estimation but not reported. Asterisks denote the level of statistical significance with *** p<0.001, ** p<0.01, * p<0.05, † p<0.1.

Appendix 1. Country Sample

Argentina, Australia, Brazil, Canada, Chile (1995), China (1997), Colombia, Costa Rica (1996), Ecuador (1998), European Union, India (1992), Indonesia (1996), Israel, Jamaica (2000), Japan, Republic of Korea, Malaysia (1995), Mexico, New Zealand (1995), Pakistan (2002), Paraguay (1999), Peru (1992), Philippines (1994), South Africa, Taiwan, Thailand (1996), Trinidad and Tobago (1997), Turkey, United States, Uruguay (1997), Venezuela (1992).

Notes: The European Union is treated as a single country. Its evolutionary enlargement of member states is considered. Unless otherwise stated in parentheses, the initial year of the respective country in our sample is 1991.

Appendix 2. Variable Summary Statistics

VARIABLE	Observations	Mean	Std. Dev.	Min	Max
$\ln(M_{ijkt})$	54,133	15.810	2.918	0	24.263
$\ln(1+t_{ijkt})$	55,649	0.080	0.087	0	0.961
$\ln\left(RER_{ijt}\right)$	59,416	-0.953	2.716	-9.698	9.698
$\ln(M\_SHARE_{ijkt})$	54,133	1.807	2.144	-13.154	4.605
$\text{before}_{ijk,T-1}$	59,920	0.031	0.173	0	1
$\text{before}_{ijk,T-2}$	59,920	0.030	0.172	0	1
$\text{before}_{ijk,T-3}$	59,920	0.029	0.168	0	1
$\text{after}_{ijk,T+1}$	59,920	0.028	0.164	0	1
$\text{after}_{ijk,T+2}$	59,920	0.026	0.159	0	1
$\text{after}_{ijk,T+3}$	59,920	0.024	0.154	0	1

Appendix 3. Variable Description and Data Source

VARIABLE	Description	Data Source
M_{ijtk}	Volume of annual imports of country i from j at the Harmonised System (HS) four-digit commodity level k in period t.	UN Comtrade (2015)
t_{ijkt}	Annual trade-weighted effectively applied tariff rate of country i towards j at the HS four-digit commodity level k in period t.	World Bank (2015)
RER_{ijt}	Annual bilateral real exchange rate of country i relative to its trading partner j ($RER_{ij} = \frac{CPI_j}{CPI_i} * \frac{NER_i}{NER_j}$) in period t where NER is annual nominal exchange rates and CPI is Consumer Price Index (See Appendix 4). To fill in missing CPI data for China for the years 1991 and 1992 the GDP deflator (annual %) was used as benchmark.	UNCTAD (2016), World Bank (2016)
M_SHARE_{ijkt}	Share of j in i's annual total imports at the HS four-digit commodity level k in period t.	Own computation based on data from UN Comtrade (2015)
$\sum_{n=T-3}^{T-1} \text{before}_{ijk,n}$	Dummy variable set consisting of three dummy variables that equal unity for the last in-force year of an AD measure that i imposes on j at the HS four-digit commodity level k (n $= T - 1$), and for its previous two years (n $= T - 2$, and n $= T - 3$, respectively), all zero otherwise.	Own computation based on data from Bown (2015)
$\sum_{m=T+1}^{T+3} \text{after}_{ijk,m}$	Dummy variable set consisting of three dummy variables that equal unity for the first full year after the revocation of a previously imposed AD measure that i imposes on j at the HS four-digit commodity level k (m $= T + 1$), and for the following two years (m $= T + 2$, and m $= T + 3$, respectively), all zero otherwise.	Own computation based on data from Bown (2015)

Appendix 4. Real Exchange Rate

The real exchange rate is generally defined as $RER_{ij} = \frac{CPI_j}{CPI_i} * \frac{NER_i}{NER_j}$ using UNCTAD data on annual nominal exchange rates (NER)[13] and Consumer Price Index (CPI) data, taken from UNCTAD statistics (primary source) and the World Bank.[14] We employ this definition for all sample importers except for the European Union.

In the case of the European Union (EU) we instead rely on the Real Effective Exchange Rate (REER) as we need to consider both the regional integration process of the EU and the relative economic importance of individual European countries.[15] We follow the methodology introduced in Schmitz et al. (2012) and Turner and Van't dack (1993) and use a double-weighted index. The Real Effective Exchange Rate (REER) of country *i* against the European Union is calculated as:

$$REER_{i,EU} = \prod_{e=1}^{N} \left(\frac{CPI_e}{CPI_i} * \frac{NER_i}{NER_e} \right)^{w_e}, e = 1,2, \dots, N \qquad (2)$$

where *N* stands for the number of countries in the reference group European Union and *e* are the EU member countries.[16] The NERs for the Eurozone are converted from fixed historical exchange rates of national currencies towards the Euro.[17]

The trade weight w_e of each country in the reference group EU is defined as:

[13] Per US dollars exchange rates are used to derive bilateral exchange rates for each country with regard to any other country.

[14] We still have to deal with missing CPI data for China for the years 1991 and 1992. In order to keep the observations the GDP deflator (annual %) was used (as benchmark) to fill the data series.

[15] European integration process:
1991-1994: Belgium, France, Germany, Italy, Luxembourg, Netherlands, Denmark, Ireland, UK, Greece, Portugal and Spain.
1995: Austria, Finland and Sweden joined.
2004: Cyprus, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovakia, Slovenia joined.
2007: Bulgaria and Romania joined.
2013: Croatia joined.

[16] *N* varies according to the European integration process.

[17] Naturally, we have to deal with some missing values for Slovenia, Slovakia, Lithuania, Latvia, Estonia, Czech Republic and Croatia for the years 1991, 1992 and 1993. In order to keep the observations, the earliest reported values were taken for previous years. Since individual CPI data only enter as a weight into the REER this is not a problem.

$$w_e = \left(\frac{m^a}{x^a + m^a}\right) w_e^m + \left(\frac{x^a}{x^a + m^a}\right) w_e^x \quad (3)$$

where w_e^m and w_e^x are the import and export weights of each European country in country i's imports and exports from the European Union.[18] m^a refers to country i's total imports from the European countries and x^a are total global exports. Following common practice of central banks, imports and exports refer to trade in manufactured goods. We again use UN Comtrade data but on the ISIC Rev.3 division 15-37 level.[19]

The import weight w_e^m is the simple share of each European country in total imports from the European Union:

$$w_e^m = m_{i,e}^a / \sum_{e=1}^{N} m_{i,e}^a \quad (4)$$

where $m_{i,e}^a$ denotes gross import flows from the European country *e* into country *i*.

Double export weights are used to capture the effect of competition faced by country *i* with respect to global producers of manufactured goods. The export weights also take into account the exports of country *i* to all trading partners *h*. The export weight is defined as follows:

$$w_e^x = \sum_{h=1}^{H} (S_{e,h} x_h), e = 1,2, \dots, N; h = 1,2, \dots, H \quad (5)$$

Where $S_{e,h}$ is the share of *e's* supply of manufactured goods in country *h*'s market and x_h is the share of each market *h* in total exports of country *i*.

$$x_h = x_{i,h}^a / \sum_{h=1}^{H} x_{i,h}^a \quad (6)$$

$$S_{e,h} = S_{e,h}^a / \sum_{h=1}^{H} S_{e,h}^a \quad (7)$$

Where $x_{i,h}^a$ is defined as gross exports from country *i* to country *h* and $S_{e,h}^a$ is gross export flows from the European country *e* to all markets *h* for $e \neq h$ and for domestic supply for

[18] The trade weights for the effective exchange rate are averaged over the EU integration period (1991-1994; 1995-2003; 2004- 2006; 2007-2012).

[19] Data on domestically produced manufactured goods is available on this classification level.

$e = h$. Data on domestically produced manufactured goods (ISIC Rev. 3, divisions 15-37) is obtained from UNCTAD statistics.