DEUTSCHE
BUNDESBANK
EUROSYSTEM

Discussion Paper

Deutsche Bundesbank
No 52/2020

Anticipation effects of protectionist U.S. trade policies

Norbert Metiu

Deutsche Bundesbank, Wilhelm-Epstein-Straße 14, 60431 Frankfurt am Main,
Postfach 10 06 02, 60006 Frankfurt am Main

Tel +49 69 9566-0

Please address all orders in writing to: Deutsche Bundesbank,
Press and Public Relations Division, at the above address or via fax +49 69 9566-3077

Internet http://www.bundesbank.de

ISBN 978–3–95729–773–0 (Printversion)
ISBN 978–3–95729–774–7 (Internetversion)

Non-technical summary

Research question

The world trading system has changed considerably over the past few years. The United States has imposed several protectionist measures that have prompted threats of retaliation. These developments have spurred new research into the aggregate effects of trade protectionism. Yet surprisingly little is known about the macroeconomic effects of U.S. trade protection on the United States' major trading partners. This paper seeks to fill that gap with new empirical evidence on the international transmission mechanism of U.S. trade policy.

Contribution

The paper contributes to a growing literature that bridges the gap between data at the microeconomic and macroeconomic level. Isolating the surprise component of protectionist policy actions from confounding factors poses several challenges. First, protectionist trade policies may be endogenous to macroeconomic factors, making it difficult to draw causal inference. Second, substantial time delays occur between the announcement of trade-protective measures and their actual implementation, which gives rise to anticipation effects. I overcome these identification issues by developing a new measure of U.S. trade policy "announcement shocks", using micro-level data on U.S. anti-dumping, countervailing duties, and safeguards for the period 1988-2015. The measure is cleansed of macroeconomic sources of endogeneity.

Results

Estimates using the new measure indicate that U.S. trade policy announcement shocks give rise to contractions in major trading partners' output and investment. Economic expectations play an important part in the transmission of an announcement shock to the real economy. Evidence shows that the effects of the shock are substantially weaker when endogenous movements in expectations following the shock are not accounted for. An analysis that quantifies the extent of newspaper coverage of U.S. trade protection suggests that media attention facilitates the propagation of protectionist shocks, consistent with an expectations channel. Note finally that the results should be interpreted with caution when considering the post-2015 events because the sample does not include U.S. trade protection measures of the Trump administration.

Nichttechnische Zusammenfassung

Forschungsfrage

Das Welthandelssystem hat sich in den letzten Jahren stark verändert. So führte die USA mehrere protektionistische Maßnahmen ein, die ihrerseits Gegenmaßnahmen der betroffenen Länder provozierten. Im Zuge dieser Entwicklungen sind die wirtschaftlichen Folgen des Handelsprotektionismus in den Fokus der Forschung gerückt. Über die makroökonomischen Auswirkungen US-amerikanischer Handelsbeschränkungen auf die wichtigsten Handelspartner der USA ist jedoch wenig bekannt. Ziel dieses Papiers ist es, diese Lücke mit neuen empirischen Erkenntnissen zum internationalen Transmissionsmechanismus der US-Handelspolitik zu schließen.

Beitrag

Das Papier trägt zu einer wachsenden Literatur bei, die die Lücke zwischen Daten auf mikroökonomischer und makroökonomischer Ebene schließt. Die Identifikation der kausalen Effekte von Protektionismus birgt mehrere Herausforderungen. Zum einen werden protektionistische Maßnahmen häufig nicht unabhängig vom heimischen Konjunkturzyklus ergriffen. Dies erschwert die Unterscheidung zwischen exogenen handelspolitischen Schocks und einer endogenen Variation protektionistischer Maßnahmen. Zum anderen können zwischen Ankündigung und Durchführung einer Maßnahme mehrere Quartale liegen, so dass bereits ein Teil der Effekte vorweggenommen werden könnte, sobald die Maßnahme angekündigt wird. Um die kausalen Effekte protektionistischer Maßnahmen abzuschätzen, wird basierend auf Mikrodaten zu US-amerikanischen Antidumpingzöllen, Ausgleichszöllen und Schutzmaßnahmen für den Zeitraum von 1988 bis 2015 ein Indikator für handelspolitische „Ankündigungsschocks“ erstellt. Der Indikator wurde um makroökonomische Einflussfaktoren bereinigt.

Ergebnisse

Schätzungen basierend auf dem neuen Indikator zeigen, dass handelspolitische Ankündigungsschocks negative Effekte auf die wirtschaftliche Aktivität und die Investitionen der wichtigsten Handelspartner der USA haben. Bei der Übertragung handelspolitischer Ankündigungsschocks auf die Realwirtschaft spielen Erwartungen eine wichtige Rolle. So gibt es Hinweise darauf, dass die Auswirkungen der Schocks deutlich geringer sind, wenn die Veränderungen der Erwartungen nicht berücksichtigt werden. Eine Analyse, die das Ausmaß der Zeitungsberichterstattung über US-amerikanischen Handelsschutzmaßnahmen quantifiziert, deutet darauf hin, dass im Einklang mit dem Erwartungskanal die Aufmerksamkeit der Medien zur Transmission protektionistischer Schocks beiträgt. Schließlich ist zu beachten, dass die Ergebnisse bei der Betrachtung der Ereignisse nach 2015 mit Vorsicht interpretiert werden sollten, da der Schätzzeitraum die US-Handelsschutzmaßnahmen der Trump-Administration nicht umfasst.

DEUTSCHE BUNDESBANK DISCUSSION PAPER NO 52/2020

Anticipation Effects of Protectionist U.S. Trade Policies*

Norbert Metiu
Deutsche Bundesbank

Abstract

This paper investigates the international spillover effects of U.S. trade protection. Using micro-level data on anti-dumping, countervailing duties, and safeguards, I develop a new measure of U.S. trade policy announcement shocks for the period 1988-2015 that is free of confounding factors. Estimates using the new measure indicate that announced, but not yet imposed, U.S. trade restrictions give rise to contractions in major trading partners' output and investment. Counterfactual results indicate that a decline in business confidence accounts for the lion's share of these anticipation effects. A narrative analysis that quantifies the extent of newspaper coverage of U.S. trade protection shows that media attention facilitates the propagation of protectionist shocks. The results are consistent with an expectations channel of trade policy.

JEL classification: E32, F13

Keywords: Anticipation Effects; Business Cycles; Announcement Shocks; Protectionism; Trade

*Correspondence: Norbert Metiu, Directorate General Financial Stability, Deutsche Bundesbank, Wilhelm-Epstein-Straße 14, 60431 Frankfurt am Main, Germany. E-mail: norbert.metiu@bundesbank.de. Acknowledgements: I am grateful to Klaus Adam, Benjamin Born, Dario Caldara, Peter Egger, Almira Enders, Niklas Gadatsch, Mathias Hoffmann, Oscar Jordà, Markus Jorra, Benedikt Kolb, Elmar Mertens, Evi Pappa, Gabriel Perez-Quiros, Esteban Prieto, Lars E.O. Svensson and Alan M. Taylor for helpful comments and discussions. I also thank Valentin Stockerl for excellent research assistance. Disclaimer: The views expressed in this paper are those of the author and do not necessarily reflect the views of the Deutsche Bundesbank or the Eurosystem.

1 Introduction

The U.S. administration has adopted a protectionist stance on trade after the 2016 presidential election. In the course of 2018, the United States imposed new tariffs on billions of dollars in imports ranging from solar panels and washing machines to steel and aluminum. U.S. tariffs have prompted threats of retaliation, leading to a lively debate on the fallout from a potential "trade war". The European Central Bank (ECB) has identified rising trade protectionism as the major source of uncertainty in global output, emphasizing the potential negative effects on general confidence.[1] Similarly, the International Monetary Fund (IMF) has warned that an escalation of trade tensions could present risks to the global economy (see International Monetary Fund, 2018). These developments have spurred new research into the aggregate effects of trade protectionism (see, e.g., Barattieri et al., 2018; Furceri et al., 2018; Caldara et al., 2020). Yet surprisingly little is known about the macroeconomic effects of U.S. trade protection on the United States' major trading partners. This paper seeks to fill that gap with new empirical evidence on the international impact of protectionist U.S. trade policies. My main contribution is to show that announced U.S. trade restrictions give rise to significant anticipation effects ahead of their implementation, consistent with an expectations channel of trade policy.

The main challenge confronting any investigation into the effects of protectionism involves the identification of exogenous trade policy shocks. Isolating the surprise component of policy actions from confounding factors is complicated by the fact that substantial time delays occur between the announcement of trade-protective measures and their actual implementation, which gives rise to anticipation effects (see, e.g., Staiger and Wolak, 1994). Moreover, protectionist trade policies may be endogenous to macroeconomic factors, making it difficult to draw causal inference (see, e.g., Bohara and Kaempfer, 1991; Bown and Crowley, 2013). I overcome these identification issues by developing a new measure of trade policy shocks that is cleansed of confounding influences. I construct the new measure from the bottom up, using micro-level data on anti-dumping, countervailing duties, and safeguards – collectively known as temporary trade barriers (TTBs). TTBs constitute the predominant policy instruments through which advanced economies have implemented trade restrictions since the early 1980s (see Bown and Crowley, 2013). The imposition of TTBs depends on the outcome of a quasi-judicial investigation by the importing country into whether a domestic industry is suffering "material injury" from "unfair" trade practices – dumped or subsidized imports – or a surge in imports. However, in practice, businesses file more requests for TTB protection in periods of weak economic performance, suggesting that TTBs are not primarily used to combat unfair trade, but rather as a means of protectionism (see, e.g., Knetter and Prusa, 2003; Bown and Crowley, 2013). The

[1] See the press conference by Mario Draghi, President of the ECB and Luis de Guindos, Vice-President of the ECB, at the ECB in Frankfurt am Main on September 13, 2018. On the economic consequences of increasing protectionism see also Deutsche Bundesbank (2017, 2020).

U.S. is internationally among the most frequent users of TTBs, turning these policies into a useful instrument for the identification of trade policy shocks.

The new shock series is derived in three steps. First, following Bown and Crowley (2013), I count the number of products imported by the U.S. between 1988:Q1 and 2015:Q4 on which U.S. government authorities initiate an investigation that may lead to the imposition of new TTBs, relying on historical information compiled by Bown (2016) from original government source documents. From the beginning of an investigation it takes several quarters before import restrictions are imposed. Staiger and Wolak (1994) show that the mere threat of future TTBs can give rise to adverse anticipation effects. Focusing on the *initiation* date of new investigations allows me to trace out these effects from the onset of the investigation. In the second step, I link the count of products subject to TTB investigations with product-level data on bilateral trade flows, producing a time series that captures the real value of U.S. imports from all trading partners facing new TTB investigations. Finally, this series is regressed on macroeconomic variables in order to isolate exogenous variation in TTB actions from confounding factors. The residual from this regression can be best thought of as a measure of U.S. trade policy "announcement shocks" that captures the impact of announced, but not yet imposed, U.S. trade restrictions.

I employ the new measure to analyze the international effects of U.S. trade protection in an open-economy vector autoregressive (VAR) framework. The analysis focuses on the euro area as a basis of comparison and provides additional evidence for Canada, China, Japan, and Mexico. My estimates indicate that U.S. trade policy announcement shocks have economically large and statistically significant effects on the euro area. An orthogonal one-standard-deviation increase in the new measure produces a sharp decline in euro area business confidence. The prospect of new trade barriers thus undermines trust in foreign firms' future business performance. As a result, euro area economic activity contracts significantly. The response of real GDP is hump shaped, displaying a drop of about 0.1 percent on impact and a peak decline of about 0.2 percent after one year compared to baseline. The peak decline in investment is about twice as large as the drop in aggregate output. One year after the initial announcement, the euro area exports about 0.5 percent less worth of goods and services in real terms relative to baseline.

Anticipated future U.S. trade barriers may affect the euro area through various channels. Negative effects may arise directly from reduced trade flows. However, even if the direct effects might be relatively small from an aggregate perspective, trade policy announcements may also generate an economic downturn indirectly, by engendering uncertainty about the future course of policy and worsening the expected economic outlook. Changes in economic expectations can be gauged from empirical measures of confidence, as shown by Barsky and Sims (2012). The systematic behavior of such measures in response to a trade policy announcement shock might thus be important in the propagation of that shock into the real economy. I quantify the contribution of confidence to the propagation of trade policy shocks in a counterfactual experiment that isolates the effects of the shock without the systematic movement of confidence. This involves tracing out the impulse responses based on the effects when the reaction of confidence is set to zero at all horizons, as proposed by Sims and Zha (2006) and Bachmann and Sims (2012). The counterfactual impulse responses of euro area aggregates are significantly weaker than the actual responses. Changes in confidence are thus central to the transmission mechanism.

The media provides a potentially important channel for the diffusion of trade-related news. I quantify the news impact of trade policy announcements by exploiting variation in press coverage of TTB cases. I perform a narrative analysis of nearly 2,000 newspaper articles published in the Financial Times between 1988 and 2015 to identify the news reports, editorials, and opinion pieces that contain a direct reference to the domestic firms filing a TTB petition, the foreign firms involved in TTB cases, or the products under investigation. The announcement shock measure is then reweighted by the amount of media attention. The correlation coefficient between the original and the news-weighted series is relatively high at 0.77, indicating that TTB cases which affect more U.S. imports tend to receive more press coverage. The estimates obtained using the news-weighted series are very similar to the baseline results, suggesting that media attention facilitates the propagation of trade policy shocks.

Estimates for the United States' other major trading partners offer additional interesting insights. The results obtained for Japan and Mexico are qualitatively and quantitatively similar to those for the euro area. Changes in confidence also contribute to the propagation of U.S. trade policy announcement shocks into these economies. A rise in U.S. trade protection leads to a significant and persistent contraction in Chinese aggregate output. However, the direct effects via trade flows seem to dominate for China as reflected in a relatively large export drop, while changes in confidence play a negligible part in the transmission. Finally, spillovers from U.S. trade policy are relatively weak for Canada.

The empirical results are subjected to various robustness checks. I show, for instance, that the baseline results carry over to impulse responses estimated using local projection and proxy-VAR methods. Moreover, they are robust to various alterations of the baseline model specification, to varying the measure of confidence, and to aggregating information on TTB cases in different ways; e.g., by limiting the product count to anti-dumping cases or to investigations that concern exclusively members of the European single market. In addition, I rule out the possibility that the initiation of U.S. TTB investigations is systematically anticipated, and I show that controlling for the presence of large shocks, the political business cycle, or non-tariff-barriers does not materially change the main conclusions.

The paper is related to several lines of research, reviewed in the next section. In Section 3, I describe the derivation of trade policy announcement shocks. Section 4 presents the empirical evidence on the international effects of U.S. trade protection. Finally, Section 5 provides concluding remarks.

2 Related Literature

The paper is most closely related to the literature on the relationship between trade protectionism and aggregate fluctuations. On the one hand, earlier work has explored the macroeconomic determinants of TTB protection; e.g., Feinberg (1989), Knetter and Prusa (2003), and Bown and Crowley (2013, 2014). These papers find evidence of a counter-cyclical relationship between macroeconomic fluctuations and requests for protection by a domestic industry. On the other hand, earlier literature has investigated the macroeconomic consequences of restrictive trade policies; e.g., Dornbusch et al. (1977), Eichengreen (1981), Krugman (1982), and Ostry and Rose (1992). The focus of this research program has shifted to trade liberalization in subsequent years, however, the literature is currently undergoing a revival. In particular, three recent studies are most similar in spirit to my paper: Barattieri et al. (2018) show that an increase in anti-dumping initiations has adverse effects on a country's domestic business cycle; Furceri et al. (2018) reach similar conclusions based on changes in applied tariff rates; and Caldara et al. (2020) document that an unexpected rise in uncertainty about higher future tariffs reduces U.S. business investment. My study differs from these papers along three key dimensions. First, it proposes a novel approach to the identification of U.S. trade policy *announcement shocks*, which accounts for macroeconomic sources of endogeneity and anticipation effects. Second, it focuses on the international dimension by providing evidence on the impact of trade policy announcement shocks on *foreign* business cycles. Third, it shows that protectionist trade policy announcements propagate through an *expectations channel* captured by a decline in business confidence, and this propagation is facilitated by news coverage of protectionist policy actions.

A large body of literature has shown that import protection through TTBs reduces international trade flows. For example, anti-dumping duties are estimated to significantly reduce exports from targeted countries by 50% to 60% on average (see Prusa, 2001; Bown and Crowley, 2007). Anti-dumping actions impose significantly higher costs on the economy than the mere tariff analysis would suggest because of additional costs created by administrative shortcomings and firm behavior influencing the outcome of anti-dumping investigations (see Bloningen and Prusa, 2003). In addition, anti-dumping investigations exert an impact on the extensive margin and can drive export suppliers entirely out of the market (see Lu et al., 2013; Besedeš and Prusa, 2017). TTB policies also generate spillover effects that affect substitute products, downstream products, and upstream intermediaries (see, e.g., Vandenbussche and Zanardi, 2010; Egger and Nelson, 2011; Erbahar and Zi, 2017). Moreover, TTBs are frequently aimed at intermediate inputs, adversely affecting production costs (see Bown, 2018a). There are also third-country effects by which the targeted country's imports from other countries are reduced (see Bown and Crowley, 2007).

The idea that TTB initiations generate anticipation effects is not new. In the micro trade literature, Staiger and Wolak (1994) show that anticipation effects substantially depress trade, reducing imports during the period of investigation by roughly half the reduction that would be expected if trade barriers were imposed from the start of the investigation. Bown (2013) and Besedeš and Prusa (2017) provide corroborating evidence. Tharakan (1995) explains anticipation effects with a kind of "harassment" suffered by defendants due to the uncertainty and expenses associated with the investigation process, even if the complaint is finally rejected. Using a structural model, Handley and Limão (2017) demonstrate that the uncertainty associated with the threat of future trade barriers can reduce consumer welfare even if it leads to no policy change. Evidence for anticipation effects of U.S. trade policy on foreign economies at the macroeconomic level is, to my best knowledge, new in the international macro literature.

My empirical results provide valuable insight for the literature studying the influence of expectations on macroeconomic fluctuations. Early theories of the business cycle attribute an important role to confidence about economic prospects in shaping actual economic activity (see Pigou, 1927; Keynes, 1936). This view has been reinvigorated in modern business cycle theory by Beaudry and Portier (2004). Subsequent work has explored the notion that news about future economic fundamentals can generate business cycle fluctuations; see, e.g., Beaudry and Portier (2006), Jaimovich and Rebelo (2009), Schmitt-Grohe and Uribe (2012), and Leduc and Sill (2013). In this context, Barsky and Sims (2012) show that surprise movements in confidence affect macroeconomic variables because they are reflective of news about future changes in expected productivity growth. Another strand of the literature shows that announcements about future policy changes affect the economy well ahead of their implementation because of anticipation effects (see, e.g., Ramey, 2011; Mertens and Ravn, 2012; Fieldhouse et al., 2018). In related work, Bachmann and Sims (2012) find that systematic movements in confidence play an important part in the transmission of government spending shocks into economic activity, particularly during recessions. My contribution is to show that the systematic behavior of confidence also matters significantly for the propagation of trade policy announcements.

From a methodological perspective, the paper adds to a growing literature that bridges the gap between data at the microeconomic and macroeconomic level. For example, Caldara et al. (2020) construct two aggregate measures of trade policy uncertainty from firm-level data on quarterly earnings calls and textual analysis of newspaper articles as initially proposed by Baker et al. (2016). In this literature, a common approach involves removing confounding factors from the aggregate series in a regression and using the residuals as a proxy for exogenous shocks in a VAR. For instance, Gilchrist and Zakrajšek (2012) combine micro data on the near universe of U.S. corporate bonds into a credit spread index that has an important predictive content for U.S. economic activity. They study the macroeconomic impact of financial shocks, derived from innovations in corporate bond spreads after removing cyclical default premia. Basset et al. (2014) use bank-level responses to the Loan Officer Opinion Survey to construct an aggregate indicator of U.S. credit supply shocks, derived by adjusting bank lending standards for macroeconomic and bank-specific factors in a regression. Miranda-Agrippino et al. (2019) use residual variation in U.S. utility patent applications as an external instrument for the identification of technology news shocks. My empirical strategy is in line with these papers.

The construction of the new shock series combines quantitative data with historical information from government source documents compiled by Bown (2016) and narrative analysis of newspaper articles. This approach shares some important similarities with a broader empirical literature that identifies exogenous policy changes from historical records, which has its origins in Friedman and Schwartz (1963). Narrative methods can be used to separate policy actions into those taken for reasons related to cyclical economic conditions and those taken for more exogenous reasons. Examples of this purely narrative approach can be found in Romer and Romer (2010) and Fieldhouse et al. (2018). In some cases, however, it is not possible to discern from the narrative record whether policy actions are taken for reasons independent of the business cycle. The standard procedure is then to remove cyclical influences from narrative data in a regression framework. Romer and Romer (2004), for instance, combine narrative information from FOMC minutes with regression analysis to derive a measure U.S. monetary policy shocks. Cloyne and Hürtgen (2016) employ this two-step identification for the United Kingdom. I also adopt this regression-based approach.

3 Derivation of Trade Policy Announcement Shocks

Identification presents a key challenge when estimating the impact of changes in trade policy. One might be tempted to simply correlate applied tariff rates with macroeconomic aggregates. This approach would not be without pitfalls, however. For one, applied tariff rates may be endogenous to economic factors, posing an identification problem (see, e.g., Bohara and Kaempfer, 1991). A different concern is that applied tariffs vary only modestly over time. As a fundamental principle, the World Trade Organization (WTO) seeks to encourage free trade. Tariffs and non-tariff trade barriers were thus steadily reduced during several rounds of multilateral trade negotiations. As a result, applied tariff rates have been relatively low and stable, on average, for the past three decades (see Figure 1).

WTO agreements do, however, allow for exceptions from the free trade principle in the form of "contingent" trade protection through TTB measures. Specifically, the WTO's Anti-Dumping Agreement allows WTO members to charge an extra import duty in order to alle-

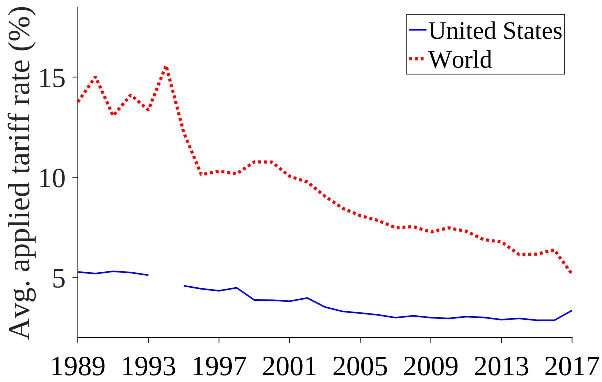


Figure 1: **Applied Tariff Rates (All Products, Mean)**

Notes: Effectively applied rates for all products subject to tariffs calculated for all traded goods (unweighted average). Annual data between 1989 and 2017. U.S. data are missing for the year 1994. Source: World Bank.

viate the injury suffered by a domestic industry from dumping, i.e., products exported at a price "less than fair value". The Agreement on Subsidies and Countervailing Measures enables WTO members to charge a countervailing duty on subsidized imports that are found to be hurting a domestic industry. Finally, the Agreement on Safeguards permits WTO members to impose safeguard tariffs on imports in order to protect domestic producers from intense import competition.

This section first provides a brief outline of contingent protection in the United States. Second, it describes three time series indicators of U.S. import protection, derived from TTB activity. Finally, residual variation in these indicators is used to infer exogenous changes in U.S. trade policy.

3.1 Contingent Protection in the United States

Historically, the U.S. has been among the countries most actively seeking contingent protection through TTB policies.[2] Prusa (2011) estimates that between 1990 and 2009 an economically important share of 4-6% of all U.S. imports were subject to trade protection, on average, through anti-dumping and countervailing duties. Moreover, I calculate that between 1988 and 2015 a cumulative share of up to approximately 4% of goods exported from the European Union (EU) to the United States were affected by U.S. TTBs.[3] There is substantial variation in

[2]Stylized facts on the use of contingent protection policies have been documented by Prusa (2011) for the United States and by Bown (2011), Bown and Crowley (2013), and European Central Bank (2013) for a broader group of countries.

[3]Based on the 5-year moving sum of the value of U.S. goods imports from EU countries affected by TTBs expressed as a fraction of all U.S. goods imports from the EU. The value of U.S. imports from EU countries subject to TTBs is computed by aggregating micro data from the United Nations Comtrade database on U.S. imports at the 6-digit, Harmonized System product level. I cumulate imports over a sliding window of five years because TTBs are typically revoked after a five year period. Aggregate data on U.S. imports come from the U.S. Census Bureau.

Anti-dumping case	Large newspaper printing presses and components thereof, whether assembled or unassembled, from Germany and Japan
U.S. petitioner	Rockwell Graphic Systems, Inc., Westmont, IL
Foreign respondent	MAN Roland Druckmaschinen AG; Koenig & Bauer-Albert AG; Mitsubishi Heavy Industries, Ltd.; Tokyo Kikai Seisakusho, Ltd.
No. products investigated (HS06 level)	10
Total trade value (1995 USD)	677,384,932
Case ID (USITC)	731-TA-736 and 731-TA-737
Case ID (ITA)	A-428-821 and A-588-837
Investigation initiated (USITC)	July 10, 1995
Investigation initiated (ITA)	July 27, 1995
Prelim. injury determination (USITC)	August 23, 1995: Affirmative
Prelim. dumping determination (ITA)	March 1, 1996: Affirmative
Final dumping determination (ITA)	July 23, 1996: Affirmative
Final decision (ITA)	September 4, 1996: Anti-dumping duty order
Final injury determination (USITC)	September 5, 1996: Affirmative
Revocation (ITA)	February 25, 2002

Table 1: **TTB Investigation Procedures: An Example from the 1990s**

Notes: Two parallel U.S. anti-dumping investigations initiated in 1995:Q3. Source: U.S. Federal Register, Bown (2016), UN Comtrade, and own calculations.

the intensity of U.S. trade protection activity across targeted countries and sectors requesting import protection. While in the 1980s about two-thirds of U.S. contingent protection cases targeted developed countries, this share fell throughout the 1990s and 2000s to about 30%. By contrast, China has emerged as the leading target of U.S. TTBs over the 2000s. Bown (2013, 2018b) documents that the U.S. steel industry is among the most frequent users of TTBs since the early 1980s, and about 20-80% of U.S. steel imports were subject to TTBs over the period between 1990 and 2017.

The imposition of TTBs is preceded by a quasi-judicial investigation conducted by the importing country.[4] Petitions for TTB protection are typically filed by firms, labor unions, or trade associations on behalf of the domestic industry, representing at least 25% of total production. On rare occasions, cases are self-initiated by the government. Anti-dumping cases are company specific, countervailing duty cases are country specific, while in principle safeguards

[4]The procedural framework of contingent protection is well documented in the literature; see, e.g., Staiger and Wolak (1994), Bagwell et al. (2009), Bown (2013), Rovegno (2013), and Besedeš and Prusa (2017). See also https://enforcement.trade.gov/intro/index.html, as well as U.S. International Trade Commission (2015) and International Trade Administration (2018a,b). The overview presented in this section is largely based on these sources.

are global. Investigations are conducted by a government authority. In the United States, TTB cases are investigated by two distinct federal agencies: the International Trade Administration (ITA) of the Department of Commerce and the United States International Trade Commission (USITC). The ITA is responsible for examining anti-competitive trade practices associated with dumping and import subsidization. The USITC is in charge of assessing injury allegations from dumping, subsidized imports, or a surge in imports.

TTB investigations are a multi-stage process, involving a preliminary and a final decision by each agency. Guidelines that describe the procedures for filing a petition requesting relief under U.S. trade laws and concerning the investigation process are provided in U.S. International Trade Commission (2015) and International Trade Administration (2018a,b). Detailed rules regulate the investigation, including criteria that specify what is meant by imports being dumped or subsidized and what constitutes a material injury. Investigations are typically completed within 12 to 18 months of initiation. Cases may not necessarily reach all stages due to withdrawal or early termination. If both agencies make affirmative final determinations, duties are applied accordingly by the U.S. Customs and Border Protection. In contrast to anti-dumping and countervailing duty cases, safeguard investigations are not required to find an unfair trade practice. Therefore, they essentially boil down to injury determination by the USITC, after which the U.S. President makes the final decision on the imposition of safeguard measures. TTBs are reviewed and either extended or lifted after a five year period.

Let us consider, for example, two parallel anti-dumping investigations from the 1990s concerning U.S. imports from Germany and Japan of large newspaper printing presses and components thereof. In the Online Appendix, I provide a detailed narrative summary of the investigations, reconstructed based on excerpts from the U.S. Federal Register. The investigation procedure can be summarized as follows (see Table 1). The cases were initiated on July 10, 1995 at the USITC and on July 27, 1995 at the ITA. The USITC made a preliminary determination on August 23, 1995 that the investigated imports were materially injuring, or threatened material injury to, the U.S. industry. The case was forwarded to the ITA which made a preliminary determination on March 1, 1996 that the products were being, or were likely to be, sold in the U.S. at less than fair value. If both agencies make affirmative preliminary determinations, "importers are required to post a bond or cash to cover an estimated amount for the duties which would be collected in the event that an [anti-dumping] or [countervailing duty] order is issued upon the completion of the investigations."[5] The final dumping decision on July 23, 1996 by the ITA and the final injury decision on September 5, 1996 by the USITC were also affirmative. The investigations were concluded by anti-dumping duty orders affecting all imports of the named products from Germany and Japan, imposed in the form of ad valorem duties – of 31% on German products and nearly 59% on Japanese products – that were subsequently

[5] See https://enforcement.trade.gov/intro/index.html; retrieved on February 20, 2019.

	Anti-dumping actions		Countervailing actions		Safeguard actions	
	1988-2015	2016-2019	1988-2015	2016-2019	1988-2015	2016-2019
Cases initiated	965	147	285	75	17	2
Trade barriers imposed	463	129	116	58	10	2

Table 2: **U.S. TTB Activity, 1988-2019**

Notes: The table reports the number of investigations initiated under U.S. trade laws concerning anti-dumping, countervailing, and safeguard actions (Cases initiated), and the number of cases in which the investigation resulted in the imposition of a temporary trade barrier (Trade barriers imposed). Source: Temporary Trade Barriers Database (Bown, 2016) for the period between 1988 and 2015, and the November 2019 issue of the WTO Report on G20 Trade Measures (World Trade Organization, 2019) for the period between January 2016 and June 2019.

revoked on February 25, 2002.

As the above example highlights, the prolonged investigation process usually takes several quarters from the date of initiation until the final implementation. Evidence suggests that the mere possibility of new trade barriers can already restrict trade ahead of the imposition of actual duties (see, e.g., Staiger and Wolak, 1994). The micro trade literature accounts for anticipation effects by focusing on the initiation of new cases instead of their final outcome (see, e.g., Feinberg, 1989; Vandenbussche and Zanardi, 2010; Egger and Nelson, 2011). The initiation date captures the relevant timing of the policy announcement, since the initiation of a new TTB investigation is publicly announced in the U.S. Federal Register.

Table 2 presents summary statistics on the initiation of U.S. contingent protection cases. The U.S. initiated 1,267 TTB investigations between 1988-2015 and another 224 cases between January 2016 and June 2019. The majority (74.6%) of investigations was conducted into dumping, and more than one half (53.2%) of the anti-dumping cases led to the imposition of extra import duties. Over the same period, 24.1% of the investigations were concerned with import subsidies, out of which 48.3% finally led to countervailing duty orders. 10 global and seven China-specific cases (1.3% of all investigations) arose under safeguard laws until the end of 2015, and two additional cases were initiated in the period thereafter.

3.2 Time Series Indicators of Import Protection

I derive three indicators of import protection based on the initiation of U.S. TTB cases. The first indicator measures the quarterly number of imported products on which a new TTB investigation is initiated in a given quarter and against which there is not already an existing TTB in place, as originally proposed by Bown and Crowley (2013). This indicator provides a more detailed picture of the intensity of TTB activity than the number of new cases because the large majority of TTB investigations concerns not one, but several products. An indicator of import protection that relies on the number of products facing new TTB initiations offers clear advantages compared to, e.g., applied tariff rates, because it is relatively free of anticipation effects and exhibits substantial time variation which can be exploited for identification purposes.

However, while the product count does well on the timing of policy announcements, it is less precise when it comes to the magnitude because it assigns equal weight to cases on products that have very different implications. Some cases refer to products with a much larger importance in terms of trade value than others. Moreover, some cases tend to receive more media attention than others. I address these issues by constructing two novel measures of import protection. In particular, I develop a second indicator which measures the trade value of products subject to new TTB initiations, obtained by linking the product count with bilateral trade data at the same disaggregation level. This is my baseline indicator of import protection. Additionally, from a narrative analysis of newspaper articles I derive a third indicator that weights the trade value of TTB initiations by the amount of media attention received by TTB cases, producing a measure of the news impact of trade policy.

3.2.1 Count of Products Subject to TTB Initiations

I compute the indicator of time-varying import protection proposed by Bown and Crowley (2013) for the United States between 1988:Q1 and 2015:Q4, based on data from the Temporary Trade Barriers Database (TTBD). The TTBD contains historical information on TTB investigations derived from government source documents by Bown (2016). Information is available on the domestic and foreign firms involved in each case, the products under investigation at the universally-defined, 6-digit Harmonized System (HS06) product level, and the relevant calendar dates: the date of initiation, the date of imposition of the preliminary measure and type of TTB imposed, the date and type of the final decision, the date of imposition of the final measure and type of measure imposed (e.g., ad valorem duty), and, if applicable, the date at which the TTB was revoked.

The sample choice is dictated by data availability. The TTBD database contains trade protection cases that stretch back to the late 1970s. However, the Harmonized System has been in place only since 1988, implying that the product count cannot be constructed at the HS06 disaggregation level before that date. The last case recorded in the database dates to late 2015. In order to capture the full breadth of the various channels, both direct and indirect, through which TTBs may potentially affect trading partners – e.g., a reduction in bilateral trade flows, third-country spillovers, a disruption of global supply chains, and anticipation effects –, the baseline count is based on the product-level initiations of new TTB investigations against all U.S. trading partners associated with all types of U.S. TTB initiations: anti-dumping, countervailing duty, and (both global and China-specific) safeguard cases. Moreover, it includes TTB initiations that did not lead to the final imposition of trade barriers. Alternative counts are considered for the sake of robustness.

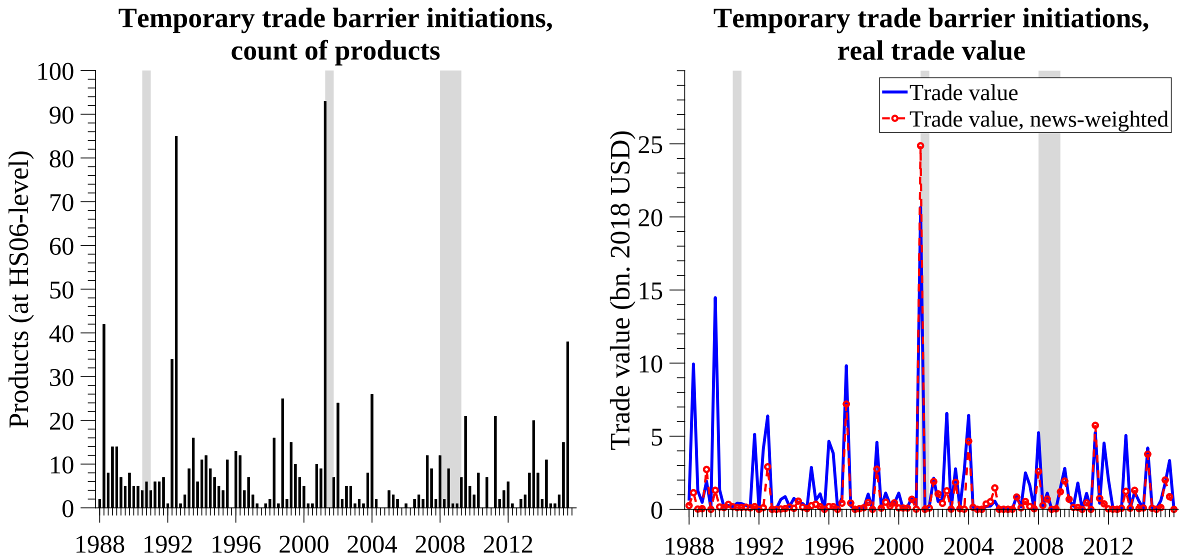


Figure 2: **Indicators of U.S. Import Protection**

Notes: Left panel: Black bars show the count of products (at the HS06 level) imported by the United States on which U.S. government agencies initiate a new temporary trade barrier (TTB) investigation in each quarter and against which there is not already an ongoing investigation or an existing TTB in place. Right panel: The solid blue line represents the real trade value (in billions of 2018 U.S. dollars) of the products subject to new TTB investigations initiated in each quarter. The red dashed line with circles corresponds to the real trade value weighted by the amount of attention devoted to each case in the Financial Times. Grey shaded regions mark recession dates published by the National Bureau of Economic Research (NBER).

To illustrate the derivation of this indicator, consider again the anti-dumping cases on German and Japanese printing presses. Investigations were conducted into alleged dumping of 10 products at the HS06 level, listed in Table A.1 of the Online Appendix. None of these products were already subject to other TTB measures at the time of initiation.[6] Moreover, all of them passed the preliminary injury determination, allowing the investigations to run their course.[7] In the third quarter of 1995 there was a single other TTB investigation initiated by U.S. authorities, concerning alleged dumping of two plastic products by U.K. producers, one of which was already facing import duties from an earlier anti-dumping order.[8] Dropping this product, the

[6]Note that parallel investigations were often conducted under more than one U.S. trade law and against multiple manufacturers exporting the same product category to the U.S. Following Bown and Crowley (2013), I do not count products if the new trade protection initiative falls into the same HS06 category as a previously imposed TTB, and I also do not count products that were subject to a simultaneous or previously imposed barrier under a different TTB measure. Moreover, also note that if the same product is subject to simultaneous investigations involving several exporting countries, the product is only counted once. Suppose, for example, that there is an anti-dumping investigation on the same product, but 12 different countries are being investigated. In this framework, that is counted as one product. I test for sensitivity with respect to these choices in robustness exercises.

[7]Note that anti-dumping and countervailing duty cases run their course only if the USITC makes an affirmative preliminary injury determination, which it must do within 45 days of initiation. Therefore, the baseline count omits initiations of anti-dumping and countervailing duty cases that were not affirmed by the USITC in its preliminary decision because these do arguably not pose a credible threat to the investigated industry. I include such cases in a robustness check.

[8]In the Online Appendix, I provide some key figures regarding this investigation.

	Product count	Trade value		
	(no. products)	(bn. 2018 US$)	(% of imports)	(bn. 2018 US$, news-weighted)
Mean	8.393	1.593	0.095	0.811
Std. dev.	13.410	2.978	0.201	2.564
Min.	0	0	0	0
Max.	93	20.652	1.301	24.869

Table 3: **Descriptive Statistics**

Notes: The table reports descriptive statistics for the count of products subject to new TTB investigations, the real trade value of products subject to new TTB investigations, the trade value of products subject to new TTB investigations in percent of aggregate U.S. imports, and the news-coverage-weighted trade value of products subject to news TTB investigations. Sample: 1988:Q1-2015:Q4. Source: TTBD, UN Comtrade, U.S. Census Bureau, FT Historical Archive, and own calculations.

total count of products subject to new TTB investigations in 1995:Q3 equals to 11 products.

Figure 2 (left panel) depicts the resulting time profile of products imported by the U.S. that were subject to TTB initiations between 1988:Q1 and 2015:Q4. Over this period, there were in total 940 products investigated at the HS06 level, averaging at more than eight products per quarter (see Table 3). Out of these, duties were finally imposed in the case of 641 products or 68% of all investigations. The standard deviation of products subject to new investigations amounts to over 13 products per quarter, suggesting that the intensity of U.S. trade protection varies considerably over time (see Table 3). Two quarters stand out with a high incidence of trade protection events. The first is associated with anti-dumping and countervailing duty investigations into steel imports initiated in 1992:Q3 against a large group of countries (85 products). The second is associated with global safeguard investigations into a range of steel products initiated in 2001:Q2 (93 products), which led to tariffs up to 30% imposed by the Bush administration in 2002 (for the historical background see Hufbauer and Goodrich, 2003; Bown, 2013).

3.2.2 Trade Value of U.S. Imports Subject to TTB Initiations

I assign time-varying weights to the product count. The weights capture the trade value of TTB initiations. Between 1988 and 2015 there were 940 products imported by the U.S. from 71 countries resulting in a total of 3,128 country-product observations that were affected by U.S. TTB investigations.[9] I derive the weights by taking the nominal (U.S. dollar) value of

[9]The investigated countries are: Argentina, Armenia, Australia, Austria, Azerbaijan, Bangladesh, Belarus, Belgium, Brazil, Canada, China, Chile, Colombia, Costa Rica, Czech Republic, Denmark, Ecuador, Egypt, Finland, France, Georgia, Germany (incl. former East and West Germany), Greece, Hong Kong, Hungary, India, Indonesia, Iraq, Ireland, Israel, Italy, Japan, Kazakhstan, Kyrgyzstan, Latvia, Lithuania, Luxembourg, Macedonia, Malaysia, Mexico, Moldova, Netherlands, New Zealand, Norway, Oman, Pakistan, Philippines, Poland, Portugal, Romania, Russia, Saudi Arabia, Singapore, Slovakia (incl. former Czechoslovakia), South Africa, South Korea, Spain, Sweden, Taiwan, Tajikistan, Thailand, Turkey, Turkmenistan, Trinidad and Tobago, Ukraine, United Arab Emirates, United Kingdom, Uzbekistan, Venezuela, Vietnam, and former Yugoslavia.

U.S. imports of each product from all exporting countries under investigation in the year the investigation was initiated. I then compute the quarterly sum of product-level import values and deflate it by the U.S. GDP-deflator to obtain real values. To construct this trade-value indicator, I compile a large micro-level data set of bilateral trade between the U.S. and all of its trading partners facing U.S. TTB initiations for the period between 1988-2015, using annual data on the trade value of U.S. imports at the HS06 level from the United Nations (UN) Comtrade Database.[10] Linking TTBD data with UN Comtrade data at the HS06 level is possible in 3,080 out of 3,128 country-product observations, producing a data coverage of 98.5%.

Considering again the above example, the nominal trade value of German and Japanese printing presses and their components imported by the U.S. in 1995 amounts to about US$ 677.4 million, while the nominal trade value of investigated U.S. imports from the U.K. equals to about US$ 15.3 million, after subtracting the value of U.K. products facing another anti-dumping duty order at the date of initiation. Thus, the total trade value of U.S. imports subject to TTB initiations in 1995:Q3 equals to US$ 692.7 million in nominal terms (1995 U.S. dollars) and US$ 1.062 billion in real terms (2018 U.S. dollars). In the Online Appendix, I provide a product-level breakdown (see Table A.1).

Figure 2 (right panel) depicts the constructed series measuring the trade value of U.S. TTB initiations between 1988:Q1 and 2015:Q4, expressed in billions of 2018 U.S. dollars. On average, over one and a half billion U.S. dollars worth of U.S. imports – or nearly 0.1 percent of total U.S. imports – face new TTB investigations each quarter, with a standard deviation of around three billion U.S. dollars per quarter (see Table 3). The 2001:Q2 steel safeguards stand out with the highest trade value of around US$ 20.65 billion, making up 1.3 percent of U.S. imports in that quarter. A notable discrepancy between the count and trade-value series is apparent in 1989:Q3. In spite of the low number of products subject to TTB initiations in that quarter (five products at the HS06 level), the trade-value indicator displays the second largest spike in the sample, exceeding US$ 14 billion. This is due to an anti-dumping case initiated in August 1989 on Canadian limousines – i.e., automobiles with a spark ignition engine of over 3,000 cc capacity (HS06 code: 870324) – that were, at the time, imported in large numbers to the United States. This case highlights the importance of accounting for the trade value of

[10] UN Comtrade data are available for the period 1991-2015. I use the 1991 values as a proxy for the years 1988-1990. Data for 9 out of 71 countries are missing. The missing countries are: Armenia, Azerbaijan, Belarus, Georgia, Kyrgyzstan, Taiwan, Tajikistan, Turkmenistan, and Uzbekistan. In some rare instances, data at the HS06 level for certain products are also not available for other countries. For the missing observations, I calculate the approximate value of product-specific U.S. imports based on trade in the missing product between the U.S. and the world from UN Comtrade scaled by the fraction of aggregate U.S. trade with the respective country relative to aggregate U.S. trade with the world obtained from the U.S. Census Bureau. For example, suppose that the value of U.S. imports of Atlantic salmon from Norway in 1992 is of interest, but that this data is not available. Suppose further that the value of U.S. imports of salmon from the world is known from the UN Comtrade database. Furthermore, according to Census Bureau data, in 1992 aggregate trade between the U.S. and Norway amounted to 0.37% of total U.S. trade with the rest of the world. The value U.S. salmon imports from the world scaled by this fraction provides an adequate approximation of the value of U.S. salmon imports from Norway in 1992.

imports in the construction of the import protection indicator.

3.2.3 Trade Value of TTB Initiations Weighted by Media Coverage

Trade policy actions potentially attract significant media attention. I quantify the extent of press coverage of TTB cases using a narrative approach. The narrative analysis relies on human readings of nearly 2,000 newspaper articles published in the Financial Times (FT) between January 1, 1988 and December 31, 2015, based on which the FT articles that contain direct references to the TTB cases documented in the TTBD database are identified. The number of mentions of each anti-dumping, countervailing duty, and safeguard case is manually recorded. This newspaper count is used to reweight the trade value of imports subject to TTB initiations by the extent of media coverage, producing a news-based measure of import protection.

I begin the narrative analysis by searching the Financial Times Historical Archive and the FT's Internet archive for articles that contain terms related to U.S. anti-dumping, countervailing duty, and safeguard cases. I find 1,727 articles containing the following combination of terms: "US" *and* "anti" *and* "dumping". Since these articles often also mention countervailing duties, I next search for "US" *and* "countervailing duties" *excluding* "anti-dumping" (with variations), resulting in 76 hits. Some of these articles report about news on TTB investigations, often mentioning several cases in the same article, while others deal with more general issues related to trade. For instance, they discuss progress on GATT/WTO negotiations. I narrow down the set of relevant FT articles by combing through the text of each article in search of the firms filing for anti-dumping and countervailing duty protection, the foreign firms under investigation, their country of origin, or the products under investigation. This information proves to be sufficient to identify those articles that contain direct references to one or more anti-dumping and countervailing duty cases. The small number of U.S. safeguard actions enables me to conduct more targeted searches by restricting the time range and using more precise search phrases to identify the articles that report on safeguard investigations.[11]

Over the sample period, I find that FT articles refer 5,756 times to anti-dumping cases, 1,325 times to countervailing duty cases, 85 times to global safeguard cases, and 19 times to China-specific safeguard cases. There is large variation in news coverage. About one third of all anti-dumping and safeguard cases and about two thirds of all countervailing duty cases were covered by the FT to varying degrees, while 705 out of 1,267 initiated cases went unreported.

[11] Global safeguards are searched within a four year window using the terms: "US" *and* "tomato" *and* "tariff", "US" *and* "brooms" *and* "tariff", "US" *and* "tomato" *and* "pepper" *and* "tariff", "US" *and* "wheat gluten" *and* "tariff", "US" *and* "lamb" *and* "tariff", "US" *and* "steel wire rod" *and* "tariff", "US" *and* "line pipe" *and* "tariff", "US" *and* "crab meat" *and* "tariff", "US" *and* "rubber thread" *and* "tariff", and "US" *and* "steel" *and* "tariff". China specific safeguards are searched using the terms: "US" *and* "China" *and* "pedestal actuators", "US" *and* "China" *and* "garment hangers", "US" *and* "China" *and* "brake drum", "US" *and* "China" *and* "waterworks fitting", "US" *and* "China" *and* "innerspring", "US" *and* "China" *and* "garment hangers", and "US" *and* "China" *and* "tyre" *and* "tariff".

The 1992 anti-dumping and countervailing duty investigations into steel imports and the global steel safeguards ten years later received most attention, at nearly 70 mentions in the FT per case.

The spectrum of media coverage ranges from news reports that inform readers about the investigation process to editorials and opinion pieces that tend to promote free trade. For instance, the anti-dumping case on German and Japanese printing presses was discussed in an article arguing that the petitioner "*[...] brought a US anti-dumping suit against its chief competitors, Mitsubishi Heavy Industries of Japan and Man Roland of Germany, after losing a large order from the Washington Post to Mitsubishi.*"[12] Another article reports on an anti-dumping investigation from 1992 into Korean dynamic random access memory (D-ram) chips:

> "*Micron Technology, a US memory chip maker, has filed a dumping complaint against South Korean semiconductor makers, accusing them of selling memory devices in the US at less than their cost of production. The anti-dumping petition, which has been filed with the US Department of Commerce and the US International Trade Commission, will prompt the US agencies to investigate the allegations of dumping [...] If dumping is proved, the US could impose dumping duties on Korean D-rams sold in the US. [...] Micron alleges that Hyundai Electronics and Goldstar Electron are worst offenders. The anti-dumping complaint applies, however, to all Korean D-ram producers, including Samsung, which last year became the world's leading producer of one megabit D-ram chips, with 14 per cent share of the world market.*" – See: "Dumping of D-rams alleged" by Louise Kehoe, in: Financial Times (London, England), Issue 31,743, p. 3, Saturday, April 25, 1992; retrieved from the Financial Times Historical Archive, 1888-2010 (Gale).

The case was formally initiated four days later on April 29, 1992 by the USITC and was concluded in May 1993 by an anti-dumping duty order that was revoked in October 2000. Meanwhile, the 2001 steel safeguard actions were criticized in an editorial piece on the grounds that they would promote protectionism:

> "*President George W. Bush's [...] threat to curb steel imports and demand for international talks to cutting overcapacity and subsidies are a cave-in to special interest lobbying and an admission that he has lost control of the trade agenda to protectionist forces in Congress. [...] As policy, Mr Bush's move is indefensible. [...] US trade restrictions would severely disrupt the world market and [...] could trigger an international spiral of protection.*" – See: "Bush's lack of steel" Editorial in: Financial

[12] See: "Rockwell to sell printing press side" in: Financial Times (London, England), Issue 32,884, p. 32, Wednesday, January 17, 1996; retrieved from the Financial Times Historical Archive, 1888-2010 (Gale).

Times (London, England), Issue 34,546, p. 20, Thursday, June 7, 2001; retrieved from the Financial Times Historical Archive, 1888-2010 (Gale).

Based on the narrative article count, I reweight the trade-value measure by the amount of media attention. The weights are specific to each case and country under investigation. A product-specific distinction at the HS06 disaggregation level is not possible in the narrative analysis. For example, the FT mentioned 'large newspaper printing presses', but no further distinction was made between, e.g., 'reel fed offset printing machines' and 'reel fed letterpress printers'. Therefore, products investigated under the same case receive equal news-coverage weight. The indicator weighted by media coverage is depicted in Figure 2 (right panel). The correlation coefficient between the trade-weighted indicator and the news-weighted series is relatively high at 0.77, indicating that TTB cases which affect a larger fraction of U.S. imports tend to receive more attention by the press.

3.3 Adjusting Import Protection for Macroeconomic Fluctuations

The existing literature suggests that TTB initiations tend to be negatively associated with movements in the business cycle because industries are inclined to file more petitions for trade protection during a slump in economic activity (see, e.g., Knetter and Prusa, 2003; Bown and Crowley, 2013). TTB initiations might thus reflect the confluence of non-cyclical policy actions and factors that are endogenous to business cycle fluctuations. This poses the following identification problem when attempting to estimate the international effects of U.S. trade protection: Suppose that a foreign economy contracts at the same time as U.S. TTB activity intensifies; the contraction abroad could be actually caused by deteriorating economic activity in the U.S. which, in turn, leads to more U.S. TTB initiations, rather than being *caused* by rising U.S. TTB protection. To rule out this common underlying cause, TTB initiations must be purged of cyclical influences.

I account explicitly for cyclical sources of endogeneity within an econometric model which is used to adjust TTB initiations for a battery of macroeconomic factors.[13] Formally, I aim to

[13] Alternatively, one could parse out the cyclical component from TTB cases based on the narrative record. However, this is complicated by several factors. First, even if firms are to some extent motivated by cyclical factors when filing a petition for TTB protection, it would be against their interest to reveal such motivation to the investigating authorities, as this would hint at rent-seeking behavior on their part. Therefore, even though TTB petitions are not public, they are highly unlikely to cite a cyclical motivation. Second, the publicly available entries in the Federal Register on TTB cases summarize the evidence submitted by petitioners substantiating their claim of unfair trade practice without any reference to cyclical developments, as illustrated in the Online Appendix for the printing press example. Regression methods are therefore better suited to remove confounding factors from TTB initiations than a purely narrative approach. This empirical strategy is in line with, e.g., Romer and Romer (2004), Basset et al. (2014), Cloyne and Hürtgen (2016), and Miranda-Agrippino et al. (2019). In principle, it is also similar to the estimation approach of factor-augmented VARs proposed by Bernanke et al. (2005), whereby the latent factors are purged of observable macroeconomic factors prior to VAR estimation.

isolate the innovations ν_t from the systematic movements in U.S. TTB initiations, TTB_t, in the following equation:

$$TTB_t = f(\boldsymbol{x}_{t-1}) + \nu_t, \tag{1}$$

where TTB_t denotes either the product count or the trade value of products subject to TTB initiations (possibly weighted by news coverage) in quarter $t = 1, ..., T$. The systematic component of TTB_t is related to macroeconomic conditions $\boldsymbol{x}_{t-1}$ via the function $f(.)$, and the term ν_t reflects the component of TTB initiations that is exogenous to economic fluctuations.

Three functional forms are considered for $f(.)$. First, if TTB_t stands for the product count, I use a negative binomial regression model following Bown and Crowley (2013) because, from a statistical perspective, the number of occurrences of an event is a count variable that takes on non-negative integer values and follows a non-normal distribution (see Cameron and Trivedi, 1986).[14] Second, if TTB_t represents the trade value or its news-weighted counterpart, the function $f(.)$ takes the form of a linear regression model, $TTB_t = \boldsymbol{x}'_{t-1}\boldsymbol{\beta} + \nu_t$, estimated by OLS. Third, in this case $f(.)$ can also take the form of a Tobit regression because the trade value series is partly continuous but censored from below at zero (for a formal exposition see Greene, 2008, Chapter 16).

Residual variation from the first-stage regression in Eq. (1) reflects U.S. trade policy actions that are orthogonal to any confounding influence from the control variables. I use the standardized regression residuals, $\xi_t = (\nu_t - E(\nu_t))/std(\nu_t))$, as a new measure of U.S. trade policy announcement shocks. It is straightforward to compute standardized residuals from the OLS and Tobit regressions. In addition, several transformations exist to compute negative binomial regression residuals that are approximately normally distributed: Pearson residuals, Anscombe (1953) residuals, and deviance-based residuals. I compute Anscombe residuals because Pierce and Schafer (1986) show that they provide a good approximation to the normal distribution in practice.[15]

Explanatory variables are selected to adequately capture cyclical fluctuations. The vector $\boldsymbol{x}_{t-1}$ contains five variables which are considered as good coincident indicators of the U.S. busi-

[14] A negative binomial regression is essentially a Poisson model with an error structure that accounts for the feature that count data tend to display pronounced over-dispersion, i.e., the conditional variance exceeds the mean. Hence, the count $TTB_t \equiv c_t$ is drawn from a Poisson distribution conditioned on explanatory variables $\boldsymbol{x}_{t-1}$ and unobserved heterogeneity η_t: $c_t|\boldsymbol{x}_{t-1}, \eta_t \sim \text{Poisson}\left(\exp(\boldsymbol{x}'_{t-1}\boldsymbol{\beta} + \eta_t)\right)$. The conditional mean of c_t admits an exponential function of the form $E(c_t|\boldsymbol{x}_{t-1}, \eta_t) \equiv \mu_t = \exp(\boldsymbol{x}'_{t-1}\boldsymbol{\beta} + \eta_t)$, where η_t is a latent term that accounts for overdispersion. The term $\exp(\eta_t)$ is gamma distributed with mean 1 and variance α (the dispersion parameter). The parameters $\boldsymbol{\beta}$ and α are jointly estimated using maximum likelihood (ML) methods (for further details see Greene, 2008, Chapter 21).

[15] Anscombe residuals are calculated as

$$\xi_t = \frac{A(c_t) - A(\widehat{\mu}_t)}{A'(\widehat{\mu}_t)\sqrt{Var(\widehat{\mu}_t)}} = \frac{\frac{3}{\hat{\alpha}}\left((1 + \hat{\alpha}c_t)^{2/3} - (1 + \hat{\alpha}\hat{\mu}_t)^{2/3}\right) + 3(c_t^{2/3} - \hat{\mu}_t^{2/3})}{2(\hat{\mu}_t + \hat{\alpha}\hat{\mu}_t^2)^{1/6}},$$

where $A(\mu) = \int_{-\infty}^{\mu} V^{-1/3}(t)dt$, $V(t)$ is the variance function, and $A'(.)$ denote the derivative of $A(.)$.

ness cycle: the log of real GDP (ΔGDP); the log of real sales in the manufacturing and trade industries (ΔSales); the unemployment rate (ΔUnemployment); the log of hours worked in the manufacturing industries (ΔHours); and the log of capacity utilization in the U.S. iron and steel industry (Capacity Utilization (Steel)). The latter is included to gauge industry-specific protectionist pressure because the U.S. steel industry is a major user of contingent protection (see Bown, 2013). Furthermore, $\boldsymbol{x}_{t-1}$ contains two variables that measure external funding constraints: the real 10-year Treasury bond yield (Real 10Y T-Bond) and a corporate bond spread index constructed by Gilchrist and Zakrajšek (2012) (GZ Credit Spread). In addition, foreign macroeconomic conditions are captured by four variables: a composite index of global real economic activity constructed by Kilian (2009) (Global Activity Index); the log of the U.S. real effective exchange rate (ΔREER); the log of real U.S. imports from the rest of the world (ΔImports); and the log of global commodity prices (ΔCommodity Prices). Variable definitions are provided in the Online Appendix. In keeping with the existing literature, the explanatory variables are expressed as year-on-year changes (with the exception of capacity utilization, the Treasury bond rate, the GZ credit spread, and Kilian's activity index), and all variables enter the regressions with a lag of one quarter to alleviate potential concerns with reverse causality. Finally, all model specifications contain a lagged dependent variable to control for potential serial dependence.

3.3.1 Measures of U.S. Trade Policy Announcement Shocks Constructed from Residual Variation in TTB Initiations

The estimation results of the first-stage regressions are shown in Table 4. The first column shows ML estimates of the negative binomial regression for counts, reported as incidence rate ratios (IRRs), i.e., exponentiated coefficients, as is standard in the literature.[16] The negative binomial estimates indicate that lower capacity utilization in the steel industry – which is a major user of contingent protection – and lower utilization of labor inputs as measured by hours worked are associated with significantly more products subject to TTB initiations. In addition, a real depreciation of the U.S. dollar and lower global commodity prices are also associated with an increased use of trade protection measures.[17] The estimated IRRs are statistically significant at the 1% level for the REER, and at the 5% level for hours, capacity utilization, and commodity

[16]The IRR represents the ratio of counts predicted by the model when an explanatory variable of interest exceeds its mean by one unit, all else equal, to the counts predicted when all variables are at their means. An IRR below one is evidence of a negative relationship (see, e.g., Bown and Crowley, 2013).

[17]The sign of the relationship between real exchange rates and filings for contingent protection is ambiguous both from a theoretical perspective and in the empirical literature (see Feinberg, 1989; Knetter and Prusa, 2003). Using data for the period 1982-1987, Feinberg (1989) finds that a real appreciation of the U.S. dollar is associated with fewer and a smaller percentage of cases. By contrast, Knetter and Prusa (2003) and Bown and Crowley (2013) find that a real appreciation of the U.S. dollar is associated with more requests for protection in samples from 1980s onwards until the Great Recession. My estimates based on data up to late 2015 conform with experience from the 1980s documented by Feinberg (1989).

	(1) Product Count (NegBin)	(2) Trade Value (OLS)	(3) Trade Value (Tobit)	(4) News-Weighted Trade Value (OLS)
Lagged Dependent	0.998	-0.147**	-0.209*	-0.105*
	(0.011)	(0.033)	(0.083)	(0.044)
ΔGDP	1.184	0.225	0.398	0.073
	(0.196)	(0.373)	(0.471)	(0.316)
ΔSales	0.998	-0.004	-0.093	0.068
	(0.115)	(0.253)	(0.320)	(0.213)
ΔWages	1.019	-0.253	-0.171	-0.298
	(0.099)	(0.233)	(0.356)	(0.279)
ΔUnemployment	1.067	-1.222	-1.017	-1.027
	(0.414)	(1.111)	(1.448)	(1.010)
ΔHours	0.735*	-0.080	-0.148	-0.287
	(0.101)	(0.306)	(0.412)	(0.236)
Capacity Utilization (Steel)	0.958*	-0.050	-0.039	-0.062
	(0.017)	(0.039)	(0.062)	(0.037)
Real 10Y T-Bond	1.157	0.426	0.447	0.155
	(0.153)	(0.310)	(0.462)	(0.190)
GZ Credit Spread	1.013	0.412	0.431	0.229
	(0.255)	(0.498)	(0.530)	(0.250)
Global Activity Index	0.999	0.001	-0.002	0.003
	(0.002)	(0.005)	(0.006)	(0.005)
ΔREER	0.923**	-0.044	-0.096	0.038
	(0.027)	(0.052)	(0.098)	(0.060)
ΔImports	1.033	0.012	-0.029	0.084
	(0.047)	(0.062)	(0.117)	(0.069)
ΔCommodity Prices	0.981*	-0.014	-0.016	0.002
	(0.008)	(0.019)	(0.018)	(0.011)
R-squared		0.066		0.084

Table 4: **Regression Estimates of U.S. TTB Protection on Macroeconomic Factors**

Notes: Column (1): Estimates of negative binomial regression with LHS: product count (exponentiated coefficients are reported). Column (2): Estimates of OLS regression with LHS: baseline indicator, i.e., the real trade value of U.S. imports subject to new TTB initiations. Column (3): Estimates of tobit regression with LHS: baseline indicator. Column (4): Estimates of OLS regression with LHS: news-weighted trade-value indicator. HAC standard errors are in parentheses. Model includes a constant term whose estimate is suppressed. Asterisks ** and * indicate statistical significance at the 1 and 5 percent levels, respectively. Sample: 1988:Q2-2015:Q4.

prices. The remaining IRRs are not statistically significant.

The second and third columns in Table 4 show OLS and Tobit estimates, respectively, for the trade-value indicator. OLS results for the news-weighted trade value are given in the last column. The results are qualitatively comparable to the negative binomial estimates in column (1), apart from evidence for some degree of first-order autocorrelation in the trade-value series. However, the relationship between trade-value indicators and macro factors is relatively weak and not statistically significant at conventional levels. Nevertheless, I opt for measuring trade policy "shocks" by the macro-adjusted series because ensuring orthogonality to macroeconomic conditions is desirable from an economic perspective. In particular, I use the standardized OLS residuals from the trade-value regression specification (2) in Table 4 as the baseline measure for U.S. trade policy announcement shocks. This series adequately captures the relevant timing of the policy announcement and its importance in terms of trade value. Moreover, it is adjusted for serial dependence and macroeconomic factors.[18]

Figure 3 shows the estimated residual shock series between 1988:Q2 and 2015:Q4. The left panel shows the standardized Anscombe residuals from a negative binomial regression that purges the TTB product count of macroeconomic factors, corresponding to specification (1) in Table 4. The middle panel depicts standardized OLS residuals ($std(\nu_t)$ = US$ 2.9 billion) from a least-squares regression of the real trade value of TTB initiations on macroeconomic factors, corresponding to specification (2) in Table 4 (the baseline shock series). Finally, the right panel shows standardized OLS residuals from a least-squares regression of the news-weighted trade value of TTB initiations on macroeconomic factors, corresponding to specification (4) in Table 4 (the news-weighted shock series). The three residual shock series are highly correlated: the correlation between the two OLS residuals equals to 0.76, while the correlation between the Anscombe residuals and trade- (news-)weighted OLS residuals is 0.63 (0.61). The 2001:Q2 shock is the largest in terms of the number, trade value, and news impact of TTB initiations.

3.3.2 Sources of the Shocks in the New Series

The objective of the regression in Equation (1) is to purge TTB initiations of cyclical factors in order to generate an exogenous proxy for trade policy shocks for the subsequent VAR analysis. However, its goal is not to find the model that explains variation in TTB initiations as well as possible by accounting for all of the determinants of TTB protection. Once I have removed the influences of fluctuations in U.S. and global economic activity from U.S. TTB actions, it is

[18] In the Online Appendix, I provide results for seven further regression specifications (see Table A.2). These differ from the baseline either in the dependent variable – by including or excluding certain types of TTB cases – or independent variables – by controlling for the political business cycle, non-tariff-barriers, or business cycle co-movement. These alternative specifications show some signs of a statistically significant link between import protection and aggregate fluctuations. I conduct VAR robustness checks using the residuals from these alternative regressions or the unpurged count or trade-value indicators as shock measures.

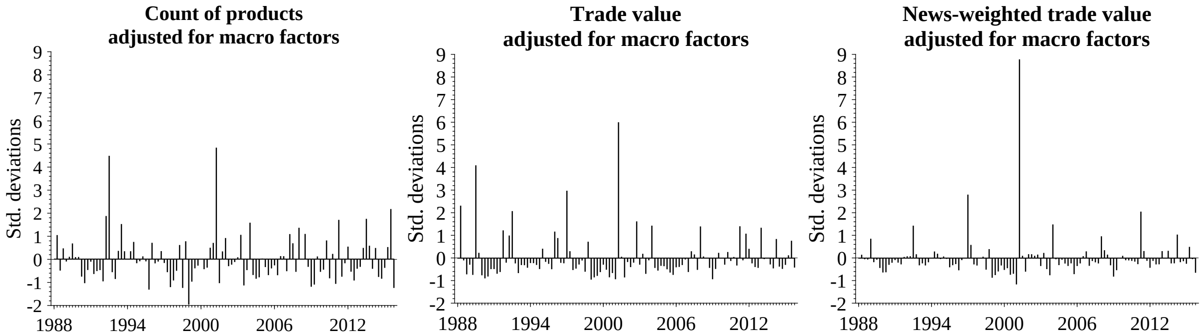


Figure 3: **Estimated Residual Shock Series**

Notes: Measures of U.S. trade policy announcement shocks constructed from residual variation in U.S. TTB initiations. Left panel: Standardized Anscombe residuals from a negative binomial regression that removes macroeconomic factors from the count of products (at the HS06 level) imported by the United States on which U.S. government agencies initiate new TTB investigations in a given quarter and against which there is not already an existing TTB in place. Middle panel: Standardized OLS residuals from a least-squares regression of the real trade value of TTB initiations on macroeconomic factors. Right panel: Standardized OLS residuals from a least-squares regression of the news-weighted trade value of TTB initiations on macroeconomic factors. Sample: 1988:Q2-2015:Q4.

desirable to leave in as much information as possible. It is precisely this residual variation that will allow me to estimate the effect of U.S. trade policy shocks on foreign economies.

Nevertheless, the estimated trade policy "shocks" do not reflect merely random movements. Instead they capture all exogenous variation in TTB initiations that arises for reasons unrelated to the business cycle. Macroeconomic conditions aside, there are at least two competing explanations for why U.S. companies file petitions for protection through TTBs (Feinberg, 1989): On the one hand, the "technical" view emphasizes the factors associated with unfair trade practices such as dumping and import subsidization; on the other hand, TTB protection cases may be promoted by rent-seeking activities on the part of petitioners. Finger et al. (1982), Grossman and Helpman (1994), and Tharakan (1995) provide theoretical support for this second – "political economy" – view. To the extent that anti-competitive trade practices by foreign firms and rent-seeking behavior by domestic firms are motivated by other than cyclical factors, both are potential sources of the estimated shocks.

4 VAR Evidence

This section introduces the VAR framework and presents the main empirical results of the paper. First, I report estimates for the euro area (EA). The EA is a useful starting point for such an investigation because it is an open economy with strong trade ties to the United States.[19] I provide additional evidence for Canada, China, Japan, and Mexico.

4.1 VAR Model

The benchmark VAR has the following representation:

$$Y_t = B_0 + B_1 Y_{t-1} + \cdots + B_p Y_{t-p} + u_t, \tag{2}$$

where $Y_t = [\xi_t, Z_t^{*'}]'$ is an $n \times 1$ vector of endogenous variables with ξ_t representing the measure of U.S. trade policy announcement shocks, measured by the real trade value of U.S. TTB initiations adjusted for macroeconomic factors, and Z_t^* denoting an $(n-1) \times 1$ vector of foreign variables. B_0 is an $n \times 1$ vector of constants, B_i are $n \times n$ coefficient matrices for $i = 1, ..., p$, where p denotes the lag order, and u_t is an $n \times 1$ vector of reduced-form innovations with $n \times n$ variance-covariance matrix $\Sigma_u = E[u_t, u_t']$.

The benchmark VAR includes the following endogenous variables for the EA, Z_t^*: (i.) log of real GDP; (ii.) log of real investment; (iii.) a composite indicator of business confidence from the OECD Business Tendency Surveys for Manufacturing, as measured by the net percent of managers reporting increased confidence in their business performance; (iv.) log of real exports; (v.) log of the Harmonised Index of Consumer Prices (HICP); (vi.) the short-term nominal interest rate; and the (vii.) log of the real exchange rate. The VAR model specifications for the other four countries are in line with the VAR for the EA, except for China, where interest rates are replaced by the log of M2 money supply because the People's Bank of China follows a monetary target (see Chen et al., 2018). Detailed variable definitions are provided in the Online Appendix. The VAR is estimated on data from 1988:Q2 to 2015:Q4. The number of lags p is set to two, in line with the Akaike criterion.

I estimate impulse responses to trade policy announcement shocks by applying the Cholesky decomposition to a recursive VAR with the shock series ordered first, followed by all other economic variables.[20] Formally, there exists a linear mapping between the reduced-form innovations u_t and an $n \times 1$ vector of mutually independent structural shocks, ε_t, given by $u_t = A\varepsilon_t$

[19] Recent data shows that the United States is the leading trading partner of the EA, accounting for 11.6% of all extra EA trade in 2017. Together with the United Kingdom and China, these three countries account for around one third of all extra EA trade (see Eurostat, 2017). In reverse, the EA is the fourth largest trading partner of the United States after China, Canada, and Mexico, accounting for 11.3% of total U.S. imports of goods in 2017, according to data from the U.S. Census Bureau; see https://www.census.gov/foreign-trade/statistics/highlights/top/top1712yr.html.

[20] My identification strategy is in line with, e.g., Ramey (2011), who also identifies shocks in a recursive VAR with the shock series ordered first. Plagborg-Møller and Wolf (2019) show that this approach is equivalent in population to the instrumental variable local projections of Stock and Watson (2018); see also the postscript to Ramey (2016). In the Online Appendix, I present impulse responses estimated by proxy-VAR and local projection methods.

with $E(\varepsilon_t \varepsilon_t') = I$, where the jth column of A represents the contemporaneous impact of the jth structural shock on the endogenous variables. Identification is achieved via recursive zero restrictions on the impact matrix A by placing the indicator ξ_t in the first position in the VAR. Orthogonal trade policy announcement shocks are then recovered through the Cholesky decomposition $\Sigma_u = AA'$ with A lower triangular. The identifying assumption implies that U.S. trade policy shocks captured by ξ_t are contemporaneously exogenous with respect to foreign business cycles. From an economic perspective, this identification strategy rests on the premise that TTB investigations are initiated based on backward-looking information, which can be rationalized by the sheer amount of data required to accompany the petition as outlined in U.S. International Trade Commission (2015) and International Trade Administration (2018a,b), potential reporting lags on the part of foreign statistical offices, as well as delays on the part of U.S. government agencies.[21]

4.2 Benchmark VAR Results for the Euro Area

The reduced-form VAR model is estimated by OLS. Figure 4 displays the impulse response functions (IRFs) of the endogenous variables to a protectionist U.S. trade policy announcement shock. The shock is estimated as an orthogonal one-standard-deviation increase in the residual shock series measuring the real trade value of U.S. TTB initiations adjusted for macroeconomic factors. The figure shows median impulse responses (black solid lines) together with the 68% (dark grey) and 90% (light grey) confidence bands based on 2,000 bootstrap draws from the estimated VAR.

A U.S. trade policy announcement shock has statistically significant macroeconomic effects on the EA. Output, investment, and exports decrease immediately, with peak responses occurring after about three quarters. Real GDP decreases by about 0.1 percent on impact relative to baseline. It reaches a peak decline of nearly 0.2 percent after three quarters, and its level is statistically indistinguishable from the baseline after two years. By comparison, a simple back-of-the-envelope calculation suggests that a shock comparable to the initiation of safeguard investigations into steel imports in 2001:Q2, which produced the largest increase of nearly six standard deviations in the trade policy indicator, leads to a three-quarter drop in EA real GDP of nearly 1.2 percent, implying that the estimated effects are economically significant. The peak decline in investment is about twice as large as the drop in aggregate output. Exports fall by around 0.5 percent relative to baseline one year after the shock. In addition, the shock causes an increase in the general price level over the very short run followed by a persistent

[21] The backward looking nature of petitions is supported by Knetter and Prusa (2003) who argue that petitioners typically assess import pricing behavior over the year prior to the filing of the case. Concerning delays on the part of the U.S. government, for instance, the initiation of an investigation can occur up to 20 days after the petition has been filed (see Staiger and Wolak, 1994; U.S. International Trade Commission, 2015). Note that, while the initiation is publicly announced, filing is not public (see Barattieri et al., 2018).

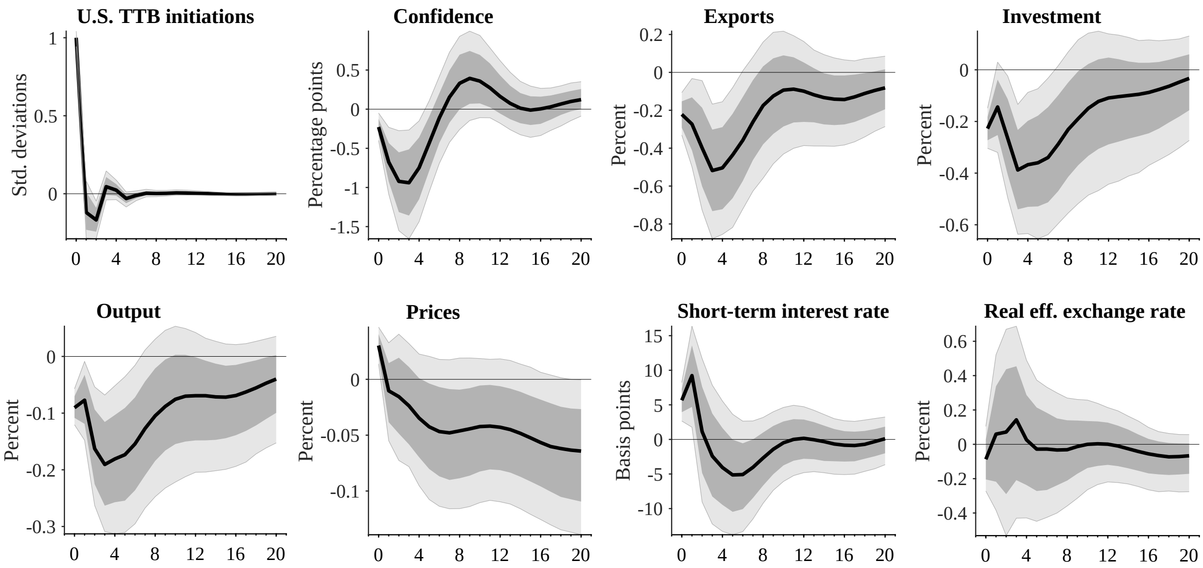


Figure 4: **Effects of a U.S. Trade Policy Announcement Shock on the Euro Area: Benchmark VAR**

Notes: Estimated impact of a one-standard-deviation U.S. trade policy announcement shock on the euro area. Impulse responses are estimated for 20 quarters. U.S. TTB initiations are measured by the real trade value of TTB initiations adjusted for macroeconomic factors. Shaded areas represent the 68% (dark grey) and 90% (light grey) confidence intervals based on 2,000 bootstrap replications. Sample: 1988:Q2-2015:Q4.

decrease. The reaction of the short-term interest rate closely tracks the price response. The real exchange rate does not display a significant reaction.

Business confidence deteriorates significantly on impact and remains below its baseline for more than a year after the initial shock. This suggests that EA businesses anticipate weaker business performance due to future trade barriers. This loss of confidence is mirrored by a decrease in investment and aggregate output, suggesting that an expectations channel of trade policy is at work. In the next section, I provide corroborating evidence that rationalizes this empirical finding.

Figure 5 shows the amount of variation in the endogenous variables explained by U.S. trade policy announcement shocks. The proportion of euro area business cycle fluctuations attributable to these shocks is relatively modest. The estimated shocks account for up to around seven percent of the variation in exports, investment, and business confidence and up to about eight percent of the variation in output at horizons associated with business cycle frequencies.

4.3 Transmission Channels: Confidence and News

There has been a long-standing view in macroeconomics that expectations and confidence about the future are relevant for understanding business cycle fluctuations (see, e.g., Pigou, 1927; Keynes, 1936; Beaudry and Portier, 2004). In line with this tradition, the contribution of ex-

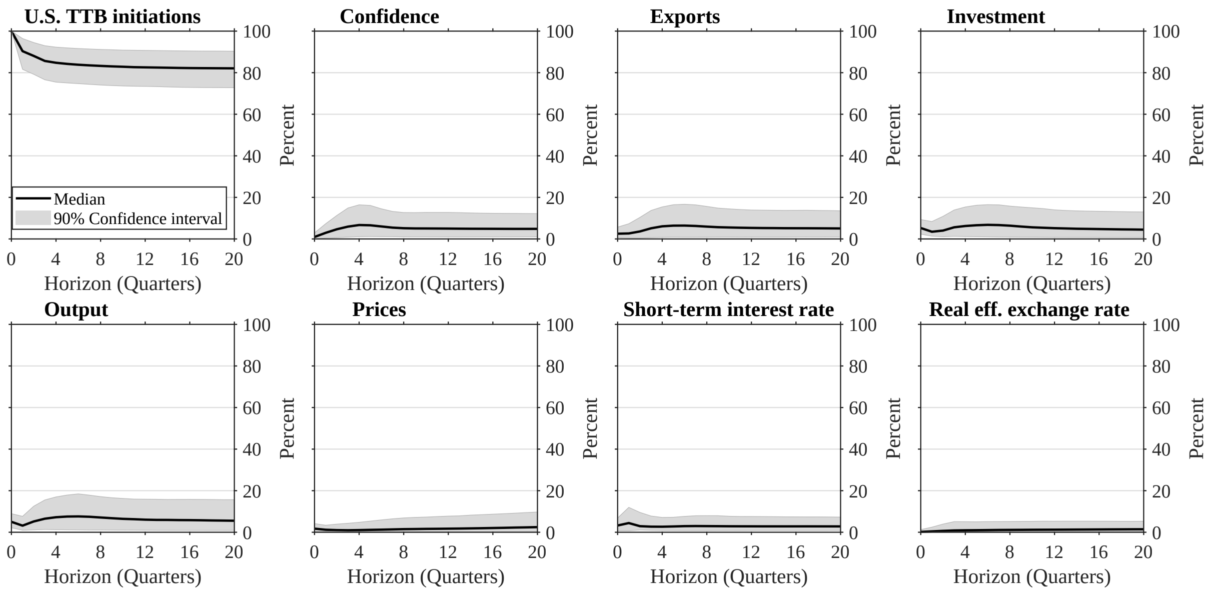


Figure 5: **Forecast Error Variance Shares Explained by U.S. Trade Policy Announcement Shocks**

Notes: The figure depicts the forecast error variance decomposition from a one-standard-deviation U.S. trade policy announcement shock on the euro area, estimated for 20 quarters. Shaded areas denote 90-percent confidence intervals based on 2,000 bootstrap replications. Sample: 1988:Q2-2015:Q4.

pectations on the part of businesses to the transmission of trade policy announcements can be quantified in a counterfactual experiment akin to those performed by Sims and Zha (2006) and Bachmann and Sims (2012). This involves constructing a scenario in which the endogenous movements of business confidence in reaction to the announcement shock are artificially switched off. Comparing the hypothetical impulse responses based on the macroeconomic effects without the endogenous response of confidence to the actual IRFs that capture the total effects allows me to measure the contribution of channels that operate through confidence to the transmission mechanism.

Figure 6 depicts the counterfactual impulse response paths, obtained using the Kalman filtering techniques of Camba-Mendez (2012). The counterfactual effects estimated while shutting down the response of confidence are statistically significant only at very short horizons and die out considerably faster than the total effects captured by the baseline IRFs. The differences between the hypothetical IRFs and the actual IRFs are economically large and statistically significant. In the case of output, for instance, the direct effects without the systematic response of confidence are only one-fourth as large as the total effects and the counterfactual point estimates lie outside of the 90% baseline confidence intervals at the trough. In conclusion, protectionist policy announcements seem to generate a downturn today chiefly because they worsen the way in which businesses perceive the future economic outlook in the face of anticipated protectionist measures.

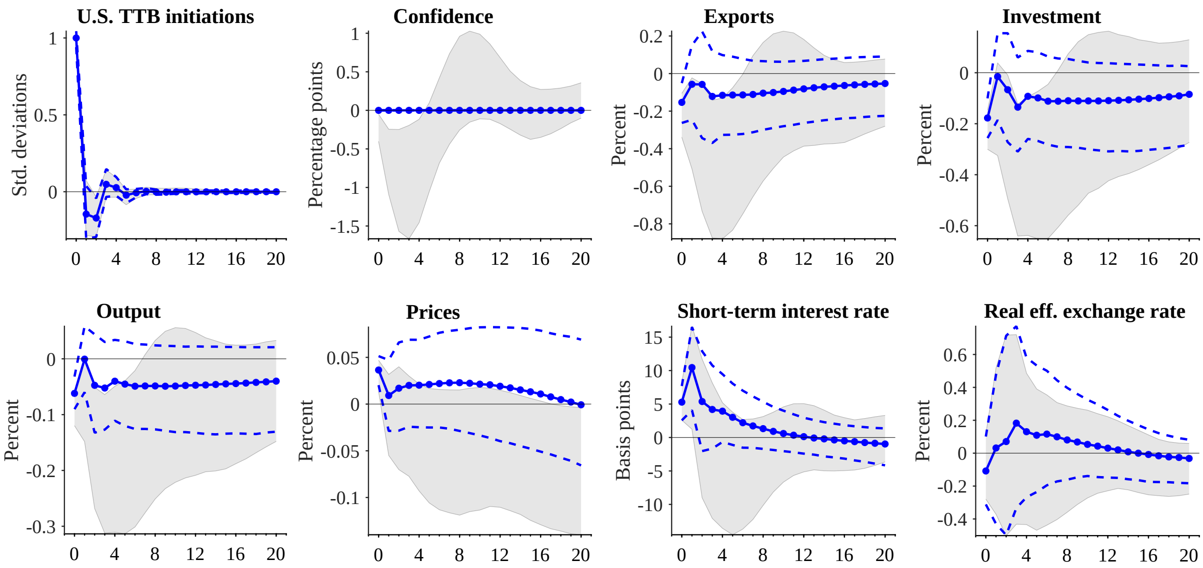


Figure 6: **Effects of a U.S. Trade Policy Announcement Shock on the Euro Area: Counterfactual Impulse Responses**

Notes: Median impulse responses (blue dotted lines) and 90% confidence bands (blue dashed lines) of euro area variables to a one-standard-deviation U.S. trade policy announcement shock based on a counterfactual scenario in which the endogenous response of business confidence is set to zero at all horizons. Impulse responses are estimated for 20 quarters. U.S. TTB initiations are measured by the real trade value of TTB initiations adjusted for macroeconomic factors. The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

I next ask whether news about trade policy actions play a part in the transmission mechanism. The median responses of EA variables computed using the news-weighted announcement shock series are displayed in Figure 7. Upon taking the news impact of trade policy into account, the macroeconomic variables for the EA display a comparable if not somewhat stronger reaction than the benchmark IRFs. The impulse responses of confidence are also in line with the benchmark VAR. The only difference is that the news-based IRFs point to a real exchange rate depreciation which is, however, economically plausible, as exporters are expected to face higher duties in the future. In sum, the results suggest that news reporting on protectionist policy actions facilitates the propagation of trade policy shocks. A potential explanation for this finding is that the media may influence public perceptions of trade policy consistent with an expectations channel.

4.4 Are TTB Initiations Systematically Anticipated?

The initiation of a new TTB investigation by U.S. government agencies represents an objective criterion to identify the policy announcement date. Nevertheless, the fact that an industry intends to file a petition might be anticipated ahead of the quarter in which the initiation takes place. To rule out the possibility that TTB announcements are *systematically* anticipated, I

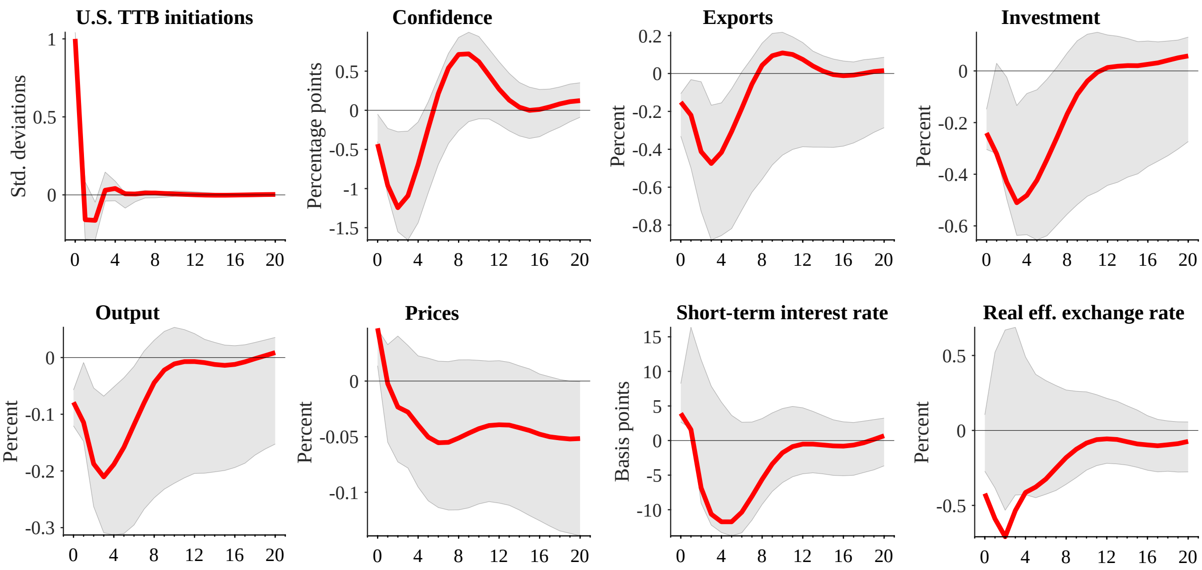


Figure 7: **Effects of a U.S. Trade Policy Announcement Shock on the Euro Area: TTB Cases Weighted by Newspaper Coverage**

Notes: Median impulse responses (red solid lines) of euro area variables to a one-standard-deviation U.S. trade policy announcement shock. Impulse responses are estimated for 20 quarters. U.S. TTB initiations are measured by the newspaper-coverage-weighted trade value of TTB initiations adjusted for macroeconomic factors. The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

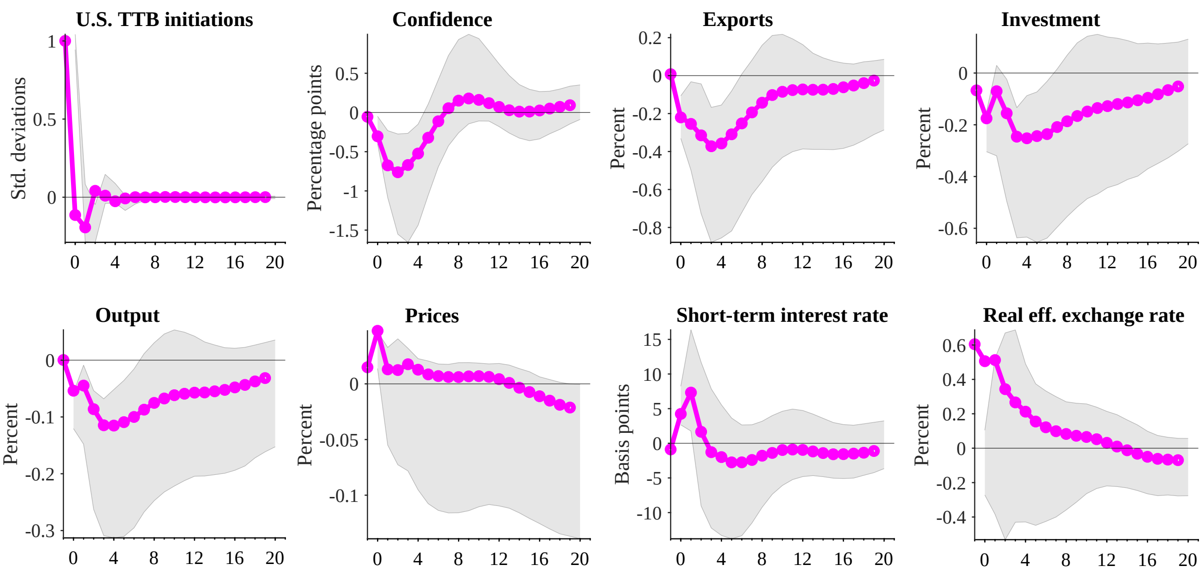


Figure 8: **Effects of a U.S. Trade Policy Announcement Shock on the Euro Area: Are U.S. TTB Initiations Systematically Anticipated?**

Notes: Median impulse responses (magenta lines with circles) of euro area variables to a one-standard-deviation U.S. trade policy announcement shock occurring in period t=-1. Impulse responses are estimated for 20 quarters. U.S. TTB initiations are measured by the real trade value of TTB initiations adjusted for macroeconomic factors, entering the VAR with one quarter lead. The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

replace the baseline indicator with its one quarter lead. If filings were anticipated in a systematic manner, this would be reflected in the impulse responses at least one quarter prior to the date of initiation. However, upon replacing the indicator with its lead, the IRFs instead resemble a staggered response pattern: the estimated effects are close to zero on impact, i.e., in the quarter preceding the initiation, and they instead materialize with a delay of one quarter (see Figure 8). The economy subsequently contracts in the quarter of initiation, broadly in line with the benchmark VAR. Hence, the shock measure adequately captures the timing of policy announcements.

4.5 Distinguishing between TTB Initiations Affecting U.S. Imports from European vs. non-European Countries

The micro data on U.S. TTBs used to derive proxies for trade policy shocks allows me to distinguish between TTB initiations affecting U.S. trade with European vs. non-European countries. To that aim, I split the trade-value indicator of import protection into the value of U.S. imports from current members of the European single market and the value of products imported from countries that are not members of the single market.[22] Computing the splitted indicators is straightforward for anti-dumping and countervailing duty initiations because they are targeted at specific countries. In addition, I also account for the share of U.S. trade with European vs. non-European countries subject to U.S. safeguards based on the country-level variation in the trade value data. Figure 9 depicts the value of U.S. goods imports from current members of the European single market and from non-European countries subject to U.S. TTB investigations initiated between 1988:Q1 and 2015:Q4. On average, the share of extra-European imports affected by U.S. TTBs is considerably larger compared to the imports from Europe facing TTB initiations. TTBs targeted at European countries are mainly concentrated in the first half of the sample.

I replace the baseline shock series in the benchmark VAR with residual shock measures derived from the two indicators displayed in Figure 9. Estimated impulse responses for the EA are depicted in Figure 10. The macroeconomic responses to protectionist U.S. trade policy shocks aimed at extra-European imports are statistically indistinguishable from the benchmark IRFs (effects are significant at the 90% level). The general picture is similar for protectionist

[22]European (EU+EFTA) countries affected by U.S. TTBs between 1988 and 2015 are: Austria, Belgium, Denmark, Finland, France, Germany (including former East and West Germany), Greece, Hungary, Ireland, Italy, Luxembourg, Netherlands, Norway, Poland, Portugal, Romania, Slovakia, Czechoslovakia, Czech Republic, Spain, Sweden, Switzerland, United Kingdom, and former Yugoslavia. Non-European countries affected by U.S. TTBs between 1988 and 2015 are: Argentina, Armenia, Australia, Azerbaijan, Bangladesh, Belarus, Brazil, Canada, Chile, Colombia, Costa Rica, Ecuador, Egypt, El Salvador, Georgia, Hong Kong, India, Iraq, Israel, Japan, Kazakhstan, Kenya, Kyrgyzstan, Macedonia, Malaysia, Mexico, Moldova, New Zealand, Oman, Pakistan, Philippines, Russia, Singapore, South Africa, South Korea, Taiwan, Tajikistan, Thailand, Trinidad and Tobago, Turkey, Turkmenistan, Ukraine, United Arab Emirates, the former USSR, Uzbekistan, Venezuela, and Vietnam.

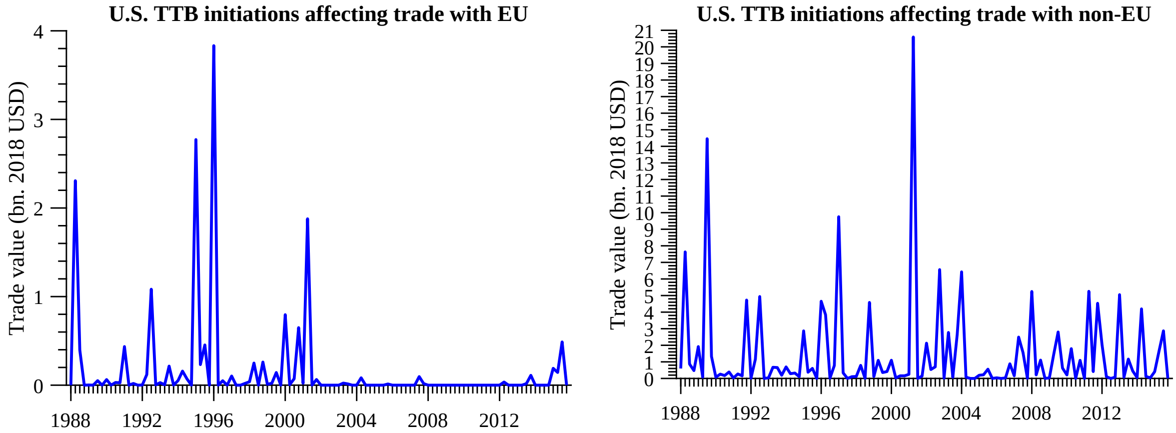


Figure 9: **U.S. TTB Initiations Affecting Trade with EU vs. Non-EU Countries**
Notes: The value of U.S. goods imports from current members of the European single market (left panel) and from non-European countries (right panel) subject to U.S. TTB investigations initiated in each quarter (in billions of 2018 U.S. dollars). Sample: 1988:Q1-2015:Q4. Source: TTBD, UN Comtrade, and own calculations.

shocks targeted at European countries. However, while EA business confidence reacts identically to both types of shocks, there are also some remarkable differences in the IRFs. Rising U.S. trade protection against extra-European imports leads to an immediate drop in the volume of EA exports, investment, and output. By contrast, the effects on euro area exports and output of TTB initiations affecting European products directly are not significant at the 90% level, and for investment they are significant only at very short horizons. These results suggest that the indirect effects of TTB protection arising from third-country spillovers dominate the direct effects on bilateral trade, underscoring the confidence effect documented above. While this finding might be surprising, it mirrors results from the micro trade literature that TTBs generate substantial spillover effects (see, e.g., Bown and Crowley, 2007; Vandenbussche and Zanardi, 2010; Erbahar and Zi, 2017). A potential explanation is that TTBs are increasingly more frequently targeting Chinese producers of intermediate inputs since the 2000s (see Bown, 2018a). This makes the EA vulnerable to protectionist shocks through international supply chains because its production strongly depends on Chinese inputs that are weakly substitutable (see European Central Bank, 2011).

4.6 Additional Robustness Exercises

I conduct several additional robustness exercises briefly summarized here; see the Online Appendix for further details. For instance, the conclusions from the recursive VAR are shown to carry over to IRFs estimated using the proxy-VAR approach by Stock and Watson (2012) and Mertens and Ravn (2013) and to IRFs obtained using the local projection method by Jordà (2005) (see Figure A.1). Moreover, I show that using a battery of alternative shock series –

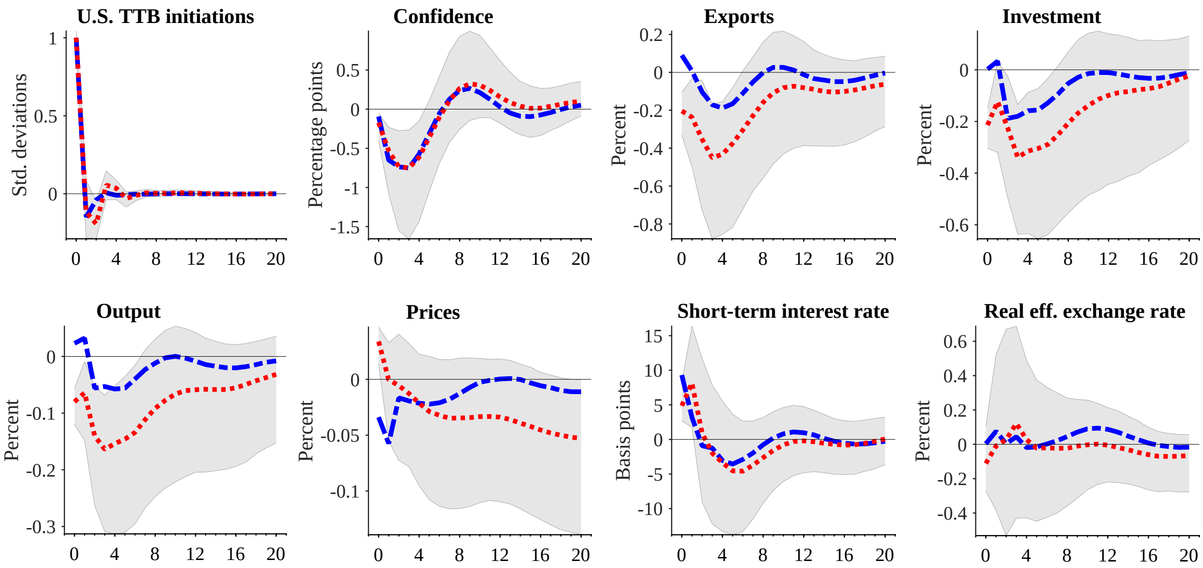


Figure 10: **Effects of a U.S. Trade Policy Announcement Shock on the Euro Area: Shocks Affecting European vs. Extra-European Imports**

Notes: Blue dashed-dotted lines represent median IRFs of EA variables to a one-standard-deviation U.S. trade policy announcement shock estimated using the shock series computed based on TTBs initiated against members of the European single market. Red dotted lines are median IRFs of euro area variables to a one-standard-deviation U.S. trade policy announcement shock estimated using the shock series computed based on TTBs initiated against non-European countries. The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

e.g., the unpurged count or trade value series, or residuals from regression specifications that control for non-tariff-barriers or the electoral cycle – leads to estimates that are statistically indistinguishable from the baseline results (see Figures A.2-A.4). In addition, I demonstrate that the omission of the two largest shock events (1989:Q3 and 2001:Q2) from the sample have no material impact on the estimated impulse responses, indicating that these events are not the sole drivers of the empirical findings (see Figure A.5). Furthermore, the macro effects are robust to adding a survey-based measure of consumer confidence to the VAR (see Figure A.5). Finally, the results are robust to estimating the VAR with more lags in the reduced form system (see Figure A.5).

4.7 Results for Canada, China, Japan, and Mexico

Figure 11 depicts the impulse responses to a U.S. trade policy announcement shock estimated for Canada, China, Japan, and Mexico. Due to data limitations, the starting date of the VAR estimations vary for these countries. The results offer some additional interesting insights relative to the EA estimates.

The estimates for Japan and Mexico are qualitatively and quantitatively similar to the effects estimated for the EA. Business confidence drops sharply for both countries after the shock,

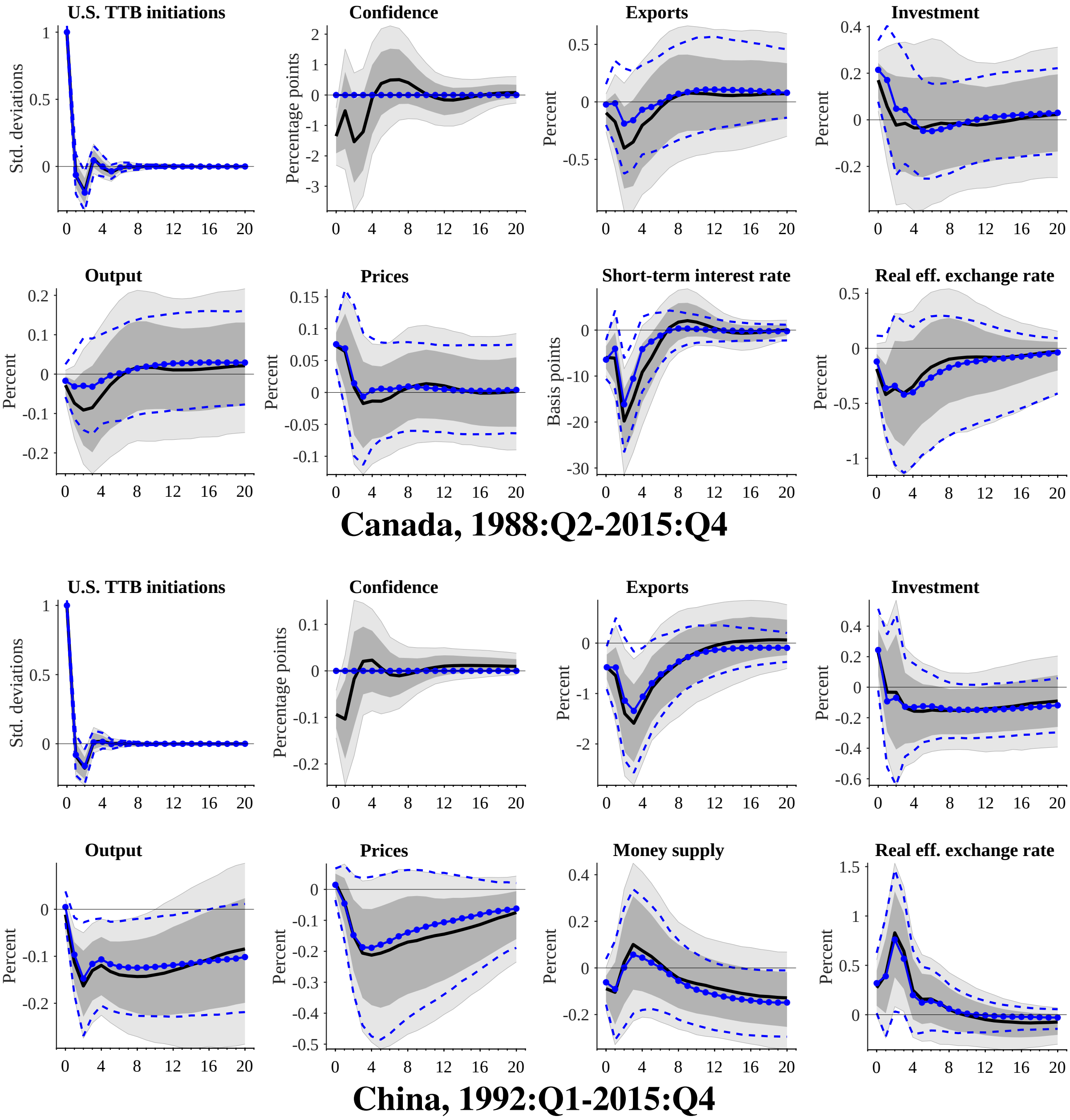
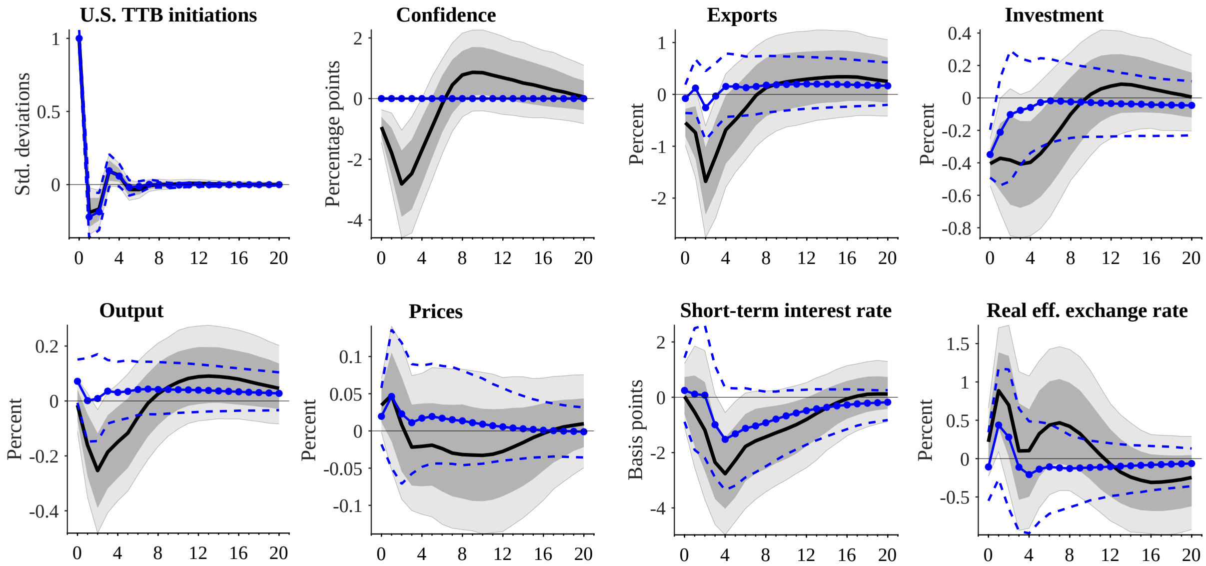
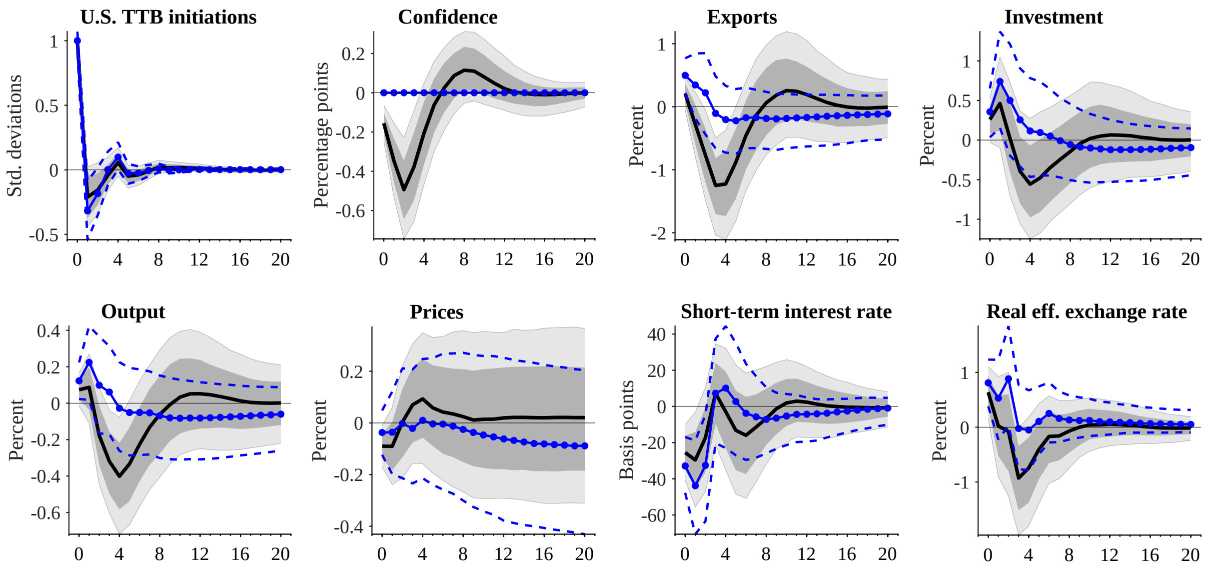


Figure 11: **Effects of a U.S. Trade Policy Announcement Shock: Canada, China, Japan, and Mexico**

Notes: Estimated impact of a one-standard-deviation U.S. trade policy announcement shock for Canada, China, Japan, and Mexico. Impulse responses are estimated for 20 quarters. U.S. TTB initiations are measured by the real trade value of TTB initiations adjusted for macroeconomic factors. Shaded areas represent the 68% (dark grey) and 90% (light grey) confidence intervals, respectively, based on 2,000 bootstrap replications. In addition, median impulse responses (blue dotted lines) and 90% confidence bands (blue dashed lines) are depicted based on a counterfactual scenario in which the endogenous response of business confidence is set to zero at all horizons. The starting date varies due to data availability.



Japan, 1994:Q1-2015:Q4



Mexico, 1998:Q1-2015:Q4

Figure 11: **Continued**

along with a significant decline in exports and aggregate output. In Japan, the real contraction is also accompanied by a relatively more pronounced drop in investment. However, all significant effects vanish once the endogenous response of business confidence is switched off. Hence, the systematic movements in confidence contribute to the propagation of trade policy announcement shocks into these two economies.

By contrast, the systematic response of confidence plays a negligible part in the transmission for China, where the effects via trade flows seem to dominate. An unexpected intensification of U.S. trade protectionism leads to a significant, relatively short, and sharp decline in Chinese exports that is nearly three times as large as the drop in EA exports after a comparable shock, with peak responses at about -1.5 percent after one year. The shock also generates a significant and persistent decrease in Chinese output. GDP in China falls by about 0.15 percent after two quarters and its response returns to baseline about ten quarters after the shock at the 90% level.

Finally, spillovers from U.S. trade policy are relatively weak in the case of Canada. Even though the Canadian IRFs display a pattern similar to the EA IRFs, the estimated effects are overwhelmingly not significant at the 90% level. A drop in the Canadian short-term interest rate seems to absorb the shock.

5 Conclusion

In this paper, I provide evidence that protectionist U.S. trade policy adversely affects foreign economies. Tracing out the international spillover effects of an exogenous U.S. trade policy shock poses a challenge for at least two reasons. First, substantial time delays occur between the initiation of new trade-protective measures and their actual implementation, which gives rise to potential anticipation effects. Second, protectionist trade policies tend to be countercyclical. The imposition of trade barriers may thus reflect the confluence of exogenous policy actions, anticipation effects, and factors that are endogenous to the business cycle. I overcome these identification issues by constructing a new measure of U.S. trade policy announcement shocks for the period 1988-2015, using micro-level data on U.S. anti-dumping, countervailing duty, and safeguard actions. I embed this measure into an otherwise standard VAR framework in order to gauge the effects of announced, but not yet imposed, U.S. trade restrictions on some of the United States' major trading partners.

I find that foreign economic activity declines immediately after protectionist U.S. trade policy announcements. Whereas direct effects through trade flows dominate for China, a counterfactual experiment shows that an endogenous deterioration of confidence in future business performance explains almost all of the macroeconomic effects for the euro area, Japan, and Mexico, suggesting that an expectations channel of trade policy plays an important part in the transmission mechanism. A narrative analysis that quantifies the extent of newspaper coverage of U.S. trade protection indicates that media attention facilitates the propagation of protectionist shocks, consistent with an expectations channel. Seeking deeper structural explanations for the relation between trade policy and expectation formation seems like a promising avenue of future research.

References

ANSCOMBE, F. (1953): "Contribution to the Discussion of H. Hotelling's Paper," *Journal of the Royal Statistical Society, Series B*, 15, 229–230.

BACHMANN, R. AND E. R. SIMS (2012): "Confidence and the Transmission of Government Spending Shocks," *Journal of Monetary Economics*, 59, 235–249.

BAGWELL, K. W., G. M. BERMANN, AND P. C. MAVROIDIS (2009): *Law and Economics of Contingent Protection in International Trade*, Cambridge University Press, Cambridge, UK.

BAKER, S., N. BLOOM, AND S. DAVIS (2016): "Measuring Economic Policy Uncertainty," *Quarterly Journal of Economics*, 131, 1593–1636.

BARATTIERI, A., M. CACCIATORE, AND F. GHIRONI (2018): "Protectionism and the Business Cycle," NBER Working Paper No. 24353.

BARSKY, R. B. AND E. R. SIMS (2012): "Information, Animal Spirits, and the Meaning of Innovations in Consumer Confidence," *American Economic Review*, 102.

BASSET, W., M. CHOSAK, J. DRISCOLL, AND E. ZAKRAJSEK (2014): "Changes in Bank Lending Standards and the Macroeconomy," *Journal of Monetary Economics*, 62, 23–40.

BAYOUMI, T., J. LEE, AND S. JAYANTHI (2005): "New Rates from New Weights," IMF Working Paper WP/05/99.

BEAUDRY, P. AND F. PORTIER (2004): "An Exploration into Pigou's Theory of Cycles," *Journal of Monetary Economics*, 51, 1183–1216.

——— (2006): "Stock Prices, News, and Economic Fluctuations," *American Economic Review*, 96, 1293–1307.

BERNANKE, B., J. BOIVIN, AND P. ELIASZ (2005): "Measuring the Effects of Monetary Policy: A Factor-Augmented Vector Autoregressive (FAVAR) Approach," *Quarterly Journal of Economics*, 120, 387–422.

Besedeš, T. and T. Prusa (2017): "The Hazardous Effects of Antidumping," *Economic Inquiry*, 55, 9–30.

Bloningen, B. A. and T. J. Prusa (2003): "The Cost of Antidumping: The Devil is in the Details," *Policy Reform*, 6, 233–245.

Bohara, A. and W. Kaempfer (1991): "A Test of Tariff Endogeneity in the United States," *American Economic Review*, 81, 952–960.

Bown, C. P. (2011): "Taking Stock of Antidumping, Safeguards and Countervailing Duties, 1990-2009," *The World Economy*, 34, 1955–1998.

——— (2013): "How Different are Safeguards from Antidumping? Evidence from US Trade Policies Toward Steel," *Review of Industrial Organization*, 42, 449–481.

——— (2016): "Temporary Trade Barriers Database," The World Bank, June 2016.

——— (2018a): "Trade Policy Toward Supply Chains after the Great Recession," *IMF Economic Review*, 66, 602–616.

——— (2018b): "US Steel is Already Highly Protected from Imports," Peterson Institute for International Economics.

Bown, C. P. and M. A. Crowley (2007): "Trade Deflection and Trade Depression," *Journal of International Economics*, 72, 176–201.

——— (2013): "Import Protection, Business Cycles, and Exchange Rates: Evidence from the Great Recession," *Journal of International Economics*, 90, 50–64.

——— (2014): "Emerging Economies, Trade Policy, and Macroeconomic Shocks," *Journal of Development Economics*, 111, 261–273.

Caldara, D., M. Iacoviello, P. Molligo, A. Prestipino, and A. Raffo (2020): "The Economic Effects of Trade Policy Uncertainty," *Journal of Monetary Economics*, 109, 38–59.

Camba-Mendez, G. (2012): "Conditional Forecasts on SVAR Models using the Kalman Filter," *Economics Letters*, 115, 376 – 378.

Cameron, A. and P. Trivedi (1986): "Econometric Models based on Count Data: Comparisons and Applications of Some Estimators and Tests," *Journal of Applied Econometrics*, 1, 29–54.

Chang, C., K. Chen, D. Waggoner, and T. Zha (2016): “Trends and Cycles in China’s Macroeconomy,” in *NBER Macroeconomics Annual*, University of Chicago Press, vol. 30, 1–84.

Chen, K., J. Ren, and T. Zha (2018): “The Nexus of Monetary Policy and Shadow Banking in China,” *American Economic Review*, 108, 3891–3936.

Cloyne, J. and P. Hürtgen (2016): “The Macroeconomic Effects of Monetary Policy: A New Measure for the United Kingdom,” *American Economic Journal: Macroeconomics*, 8, 75–102.

Deutsche Bundesbank (2017): “The Danger Posed to the Global Economy by Protectionist Tendencies,” in *Monthly Report, July 2017*, Deutsche Bundesbank, Frankfurt am Main, 77–91.

——— (2020): “Consequences of Increasing Protectionism,” in *Monthly Report, January 2020*, Deutsche Bundesbank, Frankfurt am Main, 45–66.

Dornbusch, R., S. Fischer, and P. A. Samuelson (1977): “Comparative Advantage, Trade, and Payments in a Ricardian Model with a Continuum of Goods,” *American Economic Review*, 67, 823–839.

Egger, P. and D. Nelson (2011): “How Bad is Antidumping? Evidence from Panel Data,” *Review of Economics and Statistics*, 93, 1374–1390.

Eichengreen, B. J. (1981): “A Dynamic Model of Tariffs, Output and Employment under Flexible Exchange Rates,” *Journal of International Economics*, 11, 341–359.

Enders, Z. and G. J. Mueller (2009): “On the International Transmission of Technology Shocks,” *Journal of International Economics*, 78, 45–59.

Erbahar, A. and Y. Zi (2017): “Cascading Trade Protection: Evidence from the US,” *Journal of International Economics*, 108, 274–299.

European Central Bank (2011): “Box 8: Euro Area Dependence on International Supply Chains,” in *ECB Monthly Bulletin, September 2011*, ECB, Frankfurt am Main, 87–97.

——— (2013): “Is there a Risk of a Creeping Rise in Trade Protectionism?” in *ECB Monthly Bulletin, July 2013*, ECB, Frankfurt am Main, 87–97.

Eurostat (2017): “International Trade in Goods - A Statistical Picture,” Eurostat online publication.

FAGAN, G., J. HENRY, AND R. MESTRE (2001): "An Area-Wide Model (AWM) for the Euro Area," ECB working paper No. 42.

FEINBERG, R. M. (1989): "Exchange Rates and "Unfair Trade"," *Review of Economics and Statistics*, 71, 704–707.

FIELDHOUSE, A., K. MERTENS, AND M. RAVN (2018): "The Macroeconomic Effects of Government Asset Purchases: Evidence from Postwar U.S. Housing Credit Policy," *Quarterly Journal of Economics*, 133, 1503–1560.

FINGER, J. M., H. K. HALL, AND D. R. NELSON (1982): "The Political Economy of Administered Protection," *American Economic Review*, 72, 452–466.

FRIEDMAN, M. AND A. SCHWARTZ (1963): *A Monetary History of the United States, 1867-1960*, Princeton University Press.

FURCERI, D., S. A. HANNAN, J. D. OSTRY, AND A. K. ROSE (2018): "Macroeconomic Consequences of Tariffs," Working Paper 25402, National Bureau of Economic Research.

GILCHRIST, S. AND E. ZAKRAJŠEK (2012): "Credit Spreads and Business Cycle Fluctuations," *American Economic Review*, 102, 1692–1720.

GREENE, W. H. (2008): *Econometric Analysis, 6th Edition*, Pearson Prentice Hall.

GROSSMAN, G. M. AND E. HELPMAN (1994): "Protection for Sale," *American Economic Review*, 84, 833–850.

HANDLEY, K. AND N. LIMÃO (2017): "Policy Uncertainty, Trade, and Welfare: Theory and Evidence for China and the United States," *American Economic Review*, 107, 2731–83.

HIGGIN, S. P. AND T. ZHA (2015): "China's Macroeconomic Time Series: Methods and Implications," Unpublished Manuscript, Federal Reserve Bank of Atlanta.

HUFBAUER, G. C. AND B. GOODRICH (2003): "Steel Policy: The Good, the Bad, and the Ugly," Peterson Institute for International Economics.

INTERNATIONAL MONETARY FUND (2018): "Global Financial Stability Report: A Bumpy Road Ahead," April 2018.

INTERNATIONAL TRADE ADMINISTRATION (2018a): "Guidelines for Petitions Requesting Relief under U.S. Antidumping Duty Law," International Trade Administration online publication retrieved on Oct. 25, 2018 from https://enforcement.trade.gov/petitioncounseling/pcp-tips-on-submitting.html.

——— (2018b): "Guidelines for Petitions Requesting Relief under U.S. Countervailing Duty Law," International Trade Administration online publication retrieved on Oct. 25, 2018 from https://enforcement.trade.gov/petitioncounseling/pcp-tips-on-submitting.html.

JAIMOVICH, N. AND S. REBELO (2009): "Can News about the Future Drive the Business Cycle?" *American Economic Review*, 99, 1097–1118.

JORDÀ, O. (2005): "Estimation and Inference of Impulse Responses by Local Projections," *American Economic Review*, 95(1), 161–182.

KALEMLI-OZCAN, S., E. PAPAIOANNOU, AND J.-L. PEYDRO (2013): "Financial Regulation, Financial Globalization, and the Synchronization of Economic Activity," *Journal of Finance*, 68, 1179–1228.

KEYNES, J. (1936): *The General Theory of Employment, Interest and Money*, MacMillan, New York.

KILIAN, L. (2009): "Not All Oil Price Shocks Are Alike: Disentangling Demand and Supply Shocks in the Crude Oil Market," *American Economic Review*, 99, 1053–1069.

KNETTER, M. AND T. PRUSA (2003): "Macroeconomic Factors and Antidumping Filings: Evidence from Four Countries," *Journal of International Economics*, 61, 1–17.

KRUGMAN, P. (1982): "The Macroeconomics of Protection with a Floating Exchange Rate," *Carnegie-Rochester Conference Series on Public Policy*, 16, 141–182.

LEDUC, S. AND K. SILL (2013): "Expectations and Economic Fluctuations: An Analysis Using Survey Data," *Review of Economics and Statistics*, 95, 1352–1367.

LU, Y., Z. TAO, AND Y. ZHANG (2013): "How do Exporters Respond to Antidumping Investigations?" *Journal of International Economics*, 91, 290–300.

MERTENS, K. AND M. O. RAVN (2012): "Empirical Evidence on the Aggregate Effects of Anticipated and Unanticipated US Tax Policy Shocks," *American Economic Journal: Economic Policy*, 4, 145–81.

——— (2013): "The Dynamic Effects of Personal and Corporate Income Tax Changes in the United States," *American Economic Review*, 103, 1212–47.

MIRANDA-AGRIPPINO, S., S. HACIOGLU HOKE, AND K. BLUWSTEIN (2019): "When Creativity Strikes: News Shocks and Business Cycle Fluctuations," Bank of England Staff Working Paper No. 788.

NEWEY, W. K. AND K. D. WEST (1987): "A Simple, Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica*, 55, 703–708.

OECD (2003): *Business Tendency Surveys: A Handbook*, Organisation for Economic Co-Operation and Development.

OSTRY, J. AND A. ROSE (1992): "An Empirical Evaluation of the Macroeconomic Effects of Tariffs," *Journal of International Money and Finance*, 11, 63–79.

PIERCE, D. AND D. SCHAFER (1986): "Residuals in Generalized Linear Models," *Journal of the American Statistical Association*, 81, 977–986.

PIGOU, A. (1927): *Industrial Fluctuations*, MacMillan, London.

PLAGBORG-MØLLER, M. AND C. K. WOLF (2019): "Local Projections and VARs Estimate the Same Impulse Responses," Working Paper, Princeton University.

PRUSA, T. (2001): "On the Spread and Impact of Anti-dumping," *Canadian Journal of Economics*, 34, 591–611.

PRUSA, T. J. (2011): "USA: Evolving Trends in Temporary Trade Barriers," in *The Great Recession and Import Protection: The Role of Temporary Trade Barriers*, ed. by C. P. Bown, CEPR and World Bank, London, UK, 53–83.

RAMEY, V. (2016): "Chapter 2 - Macroeconomic Shocks and Their Propagation," Elsevier, vol. 2 of *Handbook of Macroeconomics*, 71 – 162.

RAMEY, V. A. (2011): "Identifying Government Spending Shocks: It's All in the Timing," *Quarterly Journal of Economics*, 126, 1–50.

ROMER, C. D. AND D. H. ROMER (2004): "A New Measure of Monetary Shocks: Derivation and Implications," *American Economic Review*, 94, 1055–1084.

——— (2010): "The Macroeconomic Effects of Tax Changes: Estimates Based on a New Measure of Fiscal Shocks," *American Economic Review*, 100, 763–801.

ROVEGNO, L. (2013): "Trade Protection and Market Power: Evidence from US Antidumping and Countervailing Duties," *Review of World Economics*, 149, 443–476.

SCHMITT-GROHE, S. AND M. URIBE (2012): "What's News in Business Cycles," *Econometrica*, 80, 2733–2764.

SIMS, C. AND T. ZHA (2006): "Does Monetary Policy Generate Recessions?" *Macroeconomic Dynamics*, 10, 231–272.

STAIGER, R. W. AND F. A. WOLAK (1994): "Measuring Industry-Specific Protection: Antidumping in the United States," *Brookings Papers on Economic Activity. Microeconomics*, 1994, 51–118.

STOCK, J. AND WATSON (2012): "Disentangling the Channels of the 2007-2009 Recession," *Brookings Papers on Economic Activity*, Spring 2012, 81–135.

STOCK, J. H. AND M. W. WATSON (2018): "Identification and Estimation of Dynamic Causal Effects in Macroeconomics Using External Instruments," *Economic Journal*, 128, 917–948.

THARAKAN, P. K. M. (1995): "Political Economy and Contingent Protection," *Economic Journal*, 105, 1550–1564.

U.S. INTERNATIONAL TRADE COMMISSION (2015): *Publication 4540, Antidumping and Countervailing Duty Handbook, 14th Edition*, U.S. International Trade Commission Publication 4540, Washington, DC 20436.

VANDENBUSSCHE, H. AND M. ZANARDI (2010): "The Chilling Trade Effects of Antidumping Proliferation," *European Economic Review*, 54, 760–777.

WORLD TRADE ORGANIZATION (2019): "Report on G20 Trade Measures," World Trade Organization, November 21, 2019.

WU, J. AND F. XIA (2016): "Measuring the Macroeconomic Impact of Monetary Policy at the Zero Lower Bound," *Journal of Money, Credit and Banking*, 48, 253–291.

——— (2017): "Time-Varying Lower Bound of Interest Rates in Europe," Chicago Booth Research Paper No. 17-06.

Online Appendix

A Details on U.S. Anti-Dumping Investigations in 1995:Q3

U.S. trade authorities initiated three anti-dumping investigations in the third quarter of 1995. Two parallel U.S. anti-dumping investigations were conducted into "large newspaper printing presses and components thereof, whether assembled or unassembled, from Germany and Japan". A third investigation was carried out concerning "foam extruded PVC and polystyrene framing stock from the United Kingdom". Table A.1 shows the investigated products and their trade value.

HS6 Code and Product Name	Trade Value (in 1995 USD)	TTB in Place or Ongoing at Date of Initiation
Case I.: Germany and Japan		
844311: Reel fed offset printing machinery	63,071,593	No
844321: Reel fed letterpress printers except flexographic	1,667,565	No
844330: Flexographic printing machinery	62,969,378	No
844340: Gravure printing machinery	50,844,446	No
844359: Printing machinery, not elsewhere specified (e.g. other than letterpress, offset, flexographic, gravure or ink-jet printing machines)	n.a.	No
844360: Machines for uses ancillary to printing	12,289,326	No
844390: Parts of printing machinery and ancillary equipment	120,258,048	No
847149: Data processing machines; digital, automatic, presented in the form of systems	n.a.	No
847150: Units of automatic data processing machines, whether or no containing in the same housing one or two of the following types of unit: storage units, input units or output units	n.a.	No
853710: Electrical control and distribution boards, < 1kV	366,284,576	No
Case II.: United Kingdom		
392490: Plastic household, toilet articles not table, kitchen	15,282,253	No
392690: Plastic articles, not elsewhere specified	78,447,264	Yes
Total (Germany and Japan):	677,384,932	
Total (UK):	93,729,517	
Total (Germany, Japan, UK) w/o in place:	692,667,185	

Table A.1: **Products Subject to U.S. Anti-Dumping Investigations in 1995:Q3.**
Notes: Pruducts associated with U.S. anti-dumping investigations initiated in 1995:Q3. Case I.: "Large newspaper printing presses and components thereof, whether assembled or unassembled, from Germany and Japan" – Two parallel anti-dumping investigations based on petitions by Rockwell Graphic Systems, Inc. against alleged dumping by MAN Roland Druckmaschinen AG and Koenig & Bauer-Albert AG of Germany, and Mitsubishi Heavy Industries, Ltd. and Tokyo Kikai Seisakusho, Ltd. of Japan. Case II.: "Foam extruded PVC and polystyrene framing stock from the United Kingdom" – Anti-dumping investigation based on petition by Marley Mouldings, Inc. against alleged dumping by Ecoframe Plc., Robobond Ltd., and Magnolia Group Plc. of the United Kingdom. Source: U.S. Federal Register, Bown (2016), UN Comtrade, and own calculations.

In the main text of the paper, I have provided some details on the investigation procedure of the first two cases. I reconstruct below a narrative record of the investigation procedure associated with these cases, based on excerpts from the U.S. Federal Register. The third anti-dumping case was based on a petition by Marley Mouldings, Inc. against alleged dumping of plastic products by Ecoframe Plc., Robobond Ltd., and Magnolia Group Plc. of the United Kingdom. Investigation was initiated on September 18, 1995 by the USITC (case ID: 731-TA-738) and on October 6, 1995 by the ITA (case ID: A-412-817). The preliminary injury determination by the USITC was affirmative on November 8, 1995. The preliminary dumping determination by the ITA was also affirmative on May 13, 1996, and so was the ITA's final dumping determination on October 2, 1996. However, the case was subsequently withdrawn by the petitioner on October 9, 1996.

Narrative Summary of "Large Newspaper Printing Presses and Components Thereof, Whether Assembled or Unassembled, from Germany and Japan":

Federal Register, Vol. 60, No. 131, p. 35564, Jul. 10, 1995:

> **Notice of institution and scheduling of preliminary anti-dumping investigations (United States International Trade Commission (USITC)).** "The [USITC] hereby gives notice of the institution of preliminary anti-dumping investigations [...] to determine whether there is a reasonable indication that an industry in the United States is materially injured, or is threatened with material injury, or the establishment of an industry in the United States is materially retarded, by reason of imports from Germany and Japan of large newspaper printing presses and components thereof, whether assembled or unassembled [...] that are alleged to be sold in the United States at less than fair value. [...] These investigations are being instituted in response to a petition filed on June 30, 1995, by Rockwell Graphic Systems, Inc., Westmont, IL. [...] [T]he [USITC] must complete preliminary anti-dumping investigations in 45 days, or in this case by August 14, 1995. The [USITC]'s views are due at the Department of Commerce [...] by August 21, 1995."

Federal Register, Vol. 60, No. 144, p. 38546-49, Jul. 27, 1995:

> **Notice of initiation of anti-dumping duty investigations (International Trade Administration (ITA), U.S. Department of Commerce (DOC)).** "We have examined the petitions on large newspaper printing presses from Germany and Japan and have found that they meet the requirements of section 732 of the [Tariff] Act [of 1930], including the requirements concerning allegations of the material injury or threat of material injury to the domestic producers of a domestic like product by reason of the complained-of imports, allegedly sold at less than fair value. Therefore, [...] we are initiating anti-dumping duty investigations to determine whether imports of large

newspaper printing presses from Germany and Japan are being, or are likely to be, sold in the United States at less than fair value. [. . .] Unless extended, we will make our preliminary determinations by December 7, 1995."

Federal Register, Vol. 60, No. 163, p. 43816-17, Aug. 23, 1995:

Notice of preliminary determination of material injury to the U.S. industry (USITC). "On the basis of the record developed in the subject investigations, the [USITC] determines [. . .] that there is a reasonable indication that an industry in the United States is materially injured by reason of imports from Germany and Japan of large newspaper printing presses and components thereof, whether assembled or unassembled [. . .] The [USITC] transmitted its determinations in these investigations to the Secretary of Commerce on August 14, 1995."

Federal Register, Vol. 60, No. 207, p. 54841, Oct. 26, 1995:

Notice of postponement of preliminary determinations of sales at less than fair value (ITA, DOC). "On October 16, 1995, Rockwell International Corporation, the petitioner, requested that the Department postpone the preliminary determinations of these investigations by 50 days. [. . .] we are postponing the date of the preliminary determinations [. . .] until no later than January 26, 1996."

Federal Register, Vol. 61, No. 42, p. 8029-39, Mar. 1, 1996:

Notice of preliminary determination of sales at less than fair value and postponement of final determination (ITA, DOC). "[. . .] the [DOC] [. . .] has exercised its discretion to toll all deadlines for the duration of the partial shutdowns of the Federal Government from November 15 through November 21, 1995, and December 16, 1995, through January 6, 1996. Thus, all deadlines in this investigation have been extended by 28 days [. . .] We preliminarily determine that large newspaper printing presses and components thereof (LNPPs) from Japan are being, or are likely to be, sold in the United States at less than fair value (LTFV), as provided in section 733 of the [Tariff] Act [of 1930]. [. . .] The Customs Service will require a cash deposit or posting of a bond equal to the estimated amount by which the normal value exceeds the export price as shown below. [. . .] The weighted-average dumping margins are as follows: Mitsubishi Heavy Industries, Ltd.: 47.57%, Tokyo Kikai Seisakusho, Ltd.: 58.14%, All Others: 53.72%. [. . .] We preliminarily determine that [LNPPs] from Germany are being, or are likely to be, sold in the United States at [LTFV] [. . .] The weighted-average dumping margins are as follows: MAN Roland Druckmaschinen AG: 17.70%, Koenig & Bauer-Albert AG: 46.40%, All Others: 17.70%."

Federal Register, Vol. 61, No. 50, p. 10381-82, Mar. 13, 1996:

Notice of institution and scheduling of final anti-dumping investigations (USITC). "The [USITC] hereby gives notice of the institution of final anti-dumping Investigations [. . .] to determine whether an industry in the United States is materially injured or threatened with material injury, or the establishment of an industry in the United

States is materially retarded, by reason of [LTFV] imports from Germany and Japan of [LNPPs] [...] These investigations are being instituted as a result of affirmative preliminary determinations by the [DOC] that imports of large newspaper printing presses and components thereof from Germany and Japan are being sold in the United States at [LTFV] [...]"

Federal Register, Vol. 61, No. 138, p. 37283, July 17, 1996:

Notice of Commission determination to conduct a portion of the hearing in camera (USITC). "Upon request of respondents in the above-captioned final investigations, the [USITC] has unanimously determined to conduct a portion of its hearing scheduled for July 17, 1996, in camera. [...] A full discussion regarding the financial condition and related proprietary data of petitioner in these investigations can only occur if a portion of the hearing is held in camera."

Federal Register, Vol. 61, No. 172, p. 46621-24, Sep. 4, 1996:

Notice of anti-dumping duty order and amended final determination of sales at less than fair value (ITA, DOC) "[...] On August 28, 1996, the [USITC] notified the [DOC] of its final determination [...] that an industry in the United States is threatened with material injury by reason of imports of the subject merchandise from Japan. [...] the [DOC] made its final determination that [LNPPs] and components thereof from Japan are being, or are likely to be, sold in the United States at [LTFV] [...] the [DOC] will direct U.S. Customs officers to assess [...] anti-dumping duties equal to the amount by which the normal value of merchandise exceeds constructed export price of all relevant entries of LNPP from Japan. U.S. Customs officers must require, at the same time as importers would normally deposit estimated duties on this merchandise, a cash deposit equal to the estimated weighted-average anti-dumping duty margins noted below. The 'All Others' rate listed applies to all Japanese exporters of LNPP not specifically listed below. The ad valorem weighted-average dumping margins are as follows: Mitsubishi Heavy Industries, Ltd.: 62.26%, Tokyo Kikai Seisakusho, Ltd.: 56.28%, All Others: 58.69% [...] On August 28, 1996, the [USITC] notified the [DOC] of its final determination [...] that an industry in the United States is threatened with material injury by reason of imports of the subject merchandise from Germany [...] The ad valorem weighted-average dumping margins are as follows: MAN Roland Druckmaschinen AG: 30.72%, Koenig & Bauer-Albert AG: 46.40%, All Others: 3[0].72%."

Federal Register, Vol. 61, No. 180, p. 48733, Sep. 16, 1996:

Editorial correction of previously published Notice (ITA, DOC). "In notice document 96-22678 beginning on page 46623 in the issue of Wednesday, September 4, 1996, make the following correction: On page 46624, in the second column, in the table, in the second column, in the last line, '3.72' should read '30.72'."

Federal Register, Vol. 61, No. 173, p. 46824, Sep. 5, 1996:

> **Notice of final determination of material injury to the U.S. industry (USITC).** "On the basis of the record developed in the subject investigations, the [USITC] determines [...] that an industry in the United States is threatened with material injury by reason of imports from Germany and Japan of [LNPPs] and components thereof, whether assembled or unassembled, whether complete or incomplete, that have been found by the [DOC] to be sold in the United States at [LTFV] [...] The [USITC] transmitted its determinations in these investigations to the Secretary of Commerce on August 28, 1996."

Federal Register, Vol. 67, No. 37, p. 8522-23, Feb. 25, 2002:

> **Notice of final results of five-year sunset reviews and revocation of anti-dumping duty orders (ITA, DOC).** "On August 1, 2001, the [DOC] [...] initiated sunset reviews of the anti-dumping duty orders on [LNPPs] and Components Thereof, Whether Assembled or Unassembled, from Japan and Germany. One domestic interested party responded to the sunset review notice of initiation in these proceedings. However, on December 21, 2001, the domestic interested party withdrew its interest in these proceedings. Therefore, the [DOC] is revoking the anti-dumping duty orders on LNPPs from Japan and Germany. [...] Because the only domestic interested party withdrew its interest in both proceedings (see 351.218(d)(1)(i) and 351.218(e)(1)(i)(C)(1) of the Sunset Regulations), consistent with the provision of section 751(c)(3)(A) of the [Tariff] Act [of 1930], we are revoking these anti-dumping duty orders."

B Macroeconomic Data

This section provides sources and definitions of the macro variables used in the analysis. A detailed itemized description of the data is provided below.

B.1 Macro Data for Canada

Variables for Canada were retrieved from the FRED database of the Federal Reserve Bank of St. Louis and from the International Financial Statistics (IFS) of the International Monetary Fund (IMF). The variables are:

- *Business confidence.* The net percent of businesses that report increased confidence in their business performance related to present and future production of the company, order books (total and export), stocks of finished goods, tendency of future selling prices, and employment (see OECD, 2003). Quarterly, seasonally adjusted. Source: OECD Business Tendency Surveys for Manufacturing: confidence indicators: composite indicators: national indicator for Canada. Retrieved from FRED, Federal Reserve Bank of

St. Louis (code: *CANBSCICP02STSAQ*). Data is not available for the period 1988:Q1-1998:Q1. For this period, a composite indicator of consumer confidence from the OECD Consumer Opinion Surveys is used (FRED code: *CSCICP03CAM665S*).

- *Effective exchange rate (real).* Trade-weighted real effective exchange rate (REER) index. An increase in REER implies that exports become more expensive and imports become cheaper; therefore, an increase indicates a loss in trade competitiveness. Details on the construction of the IMF's REER indices can be found in Bayoumi et al. (2005). Source: IMF IFS.

- *Exports (real).* Exports of goods and services; index (2015=100). Quarterly, seasonally adjusted. Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *NAEXKP06CAQ661S*).

- *GDP (real).* Current price gross domestic product (in Canadian dollars), divided by the GDP implicit price deflator. Quarterly, seasonally adjusted. Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *CANGDPNQDSMEI* / *CANGDPDEFQISMEI*).

- *Investment (real).* Gross fixed capital formation; index (2015=100). Quarterly, seasonally adjusted. Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *NAEXKP04CAQ661S*).

- *Price level.* Consumer price index: total, all items; index (2015=100). Quarterly, adjusted for seasonal effects using the Census X-12 method (own calculations). Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *CPALCY01CAQ661N*).

- *Short-term nominal interest rate.* Yield on Government of Canada Treasury Bills; percent per annum. Quarterly average, not seasonally adjusted. Source: IMF IFS. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *INTGSTCAM193N*).

B.2 Macro Data for China

Variables for China were retrieved from the FRED database of the St. Louis Fed, the dataset compiled by Chang et al. (2016), and the IMF's IFS database. For the detailed methodology of constructing the Chang et al. dataset, see Higgin and Zha (2015). The variables are:

- *Business confidence.* The net percent of businesses that report increased confidence in their business performance related to present and future production of the company, order books (total and export), stocks of finished goods, tendency of future selling prices, and employment (see OECD, 2003); normalized (normal=100). Quarterly, seasonally adjusted. Source: OECD, Business Tendency Surveys for Manufacturing: confidence indicators: composite indicators: OECD indicator for China. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *BSCICP03CNM665S*).

- *Effective exchange rate (real).* Trade-weighted real effective exchange rate (REER) index. An increase in REER implies that exports become more expensive and imports become cheaper; therefore, an increase indicates a loss in trade competitiveness. Details on the construction of the IMF's REER indices can be found in Bayoumi et al. (2005). Source: IMF IFS.

- *Exports (real).* Exports of goods reported by the Chinese customs (RMB million), deflated by the CPI. Quarterly, seasonally adjusted. Source: Chang et al. (2016) (code: *NominalExportsGoods / CPI*).
- *GDP (real).* GDP by expenditure (RMB billion), divided by the GDP implicit price deflator. Quarterly, seasonally adjusted. Source: Chang et al. (2016) (code: *NominalGDP / GDPDeflator*).
- *Investment (real).* Gross fixed capital formation with no inventories (RMB billion), deflated by the price index for gross fixed capital formation. Quarterly, seasonally adjusted. Source: Chang et al. (2016) (code: *NominalGFCF / GFCFPriceIndex*).
- *Money supply (real).* M2 (RMB million), deflated by the CPI. Quarterly, seasonally adjusted. Source: Chang et al. (2016) (code: *M2 / CPI*).
- *Price level.* Consumer price index. Quarterly, seasonally adjusted. Source: Chang et al. (2016) (code: *CPI*).

B.3 Macro Data for the Euro Area

The majority of euro area variables were retrieved from the Area-Wide Model (AWM) Database (17th update) developed by Fagan et al. (2001). They construct synthetic euro area variables by taking weighted averages for Austria, Belgium, Germany, Finland, France, Ireland, Italy, Luxembourg, Netherlands, Portugal, and Spain. The weights used are constant GDP at market prices (PPP) for the EU 11 for 1995. The EU 11 weights are reported on page 53 in Fagan et al. (2001). If not all country series are available then the weights are re-scaled from the original EU 11 weights. Further sources of the euro area data are: the ECB Statistical Data Warehouse (SDW), the IMF IFS, and Wu and Xia (2017). The variables are:

- *Business confidence.* The net percent of businesses that report increased confidence in their business performance related to present and future production of the company, order books (total and export), stocks of finished goods, tendency of future selling prices, and employment (see OECD, 2003). Quarterly, seasonally adjusted. Source: OECD Business Tendency Surveys for Manufacturing: confidence Indicators: composite Indicators: European Commission and national indicators for the euro area. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *BSCICP02EZM460S*).
- *Consumer confidence.* The net percent of households reporting increased consumer confidence based on balance of positive over negative answers to the following four questions with five answer alternatives to each question (a lot better, a little better, the same, a little worse, a lot worse): (1) Expected change in financial situation of household over the next 12 months; (2) Expected change in general economic situation over next 12 months;

(3) Expected change in unemployment over the next 12 months; (4) Expected change in savings of household over next 12 months. Quarterly, seasonally Aajusted. Source: OECD Consumer Opinion Surveys: confidence indicators: composite indicators: European Commission and national indicators for the euro area. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *CSCICP02EZM460S*).

- *Effective exchange rate (real).* Trade-weighted real effective exchange rate (REER) index. An increase in REER implies that exports become more expensive and imports become cheaper; therefore, an increase indicates a loss in trade competitiveness. Details on the construction of the IMF's REER indices can be found in Bayoumi et al. (2005). Source: IMF IFS.

- *Exports (real).* Exports of goods and services (millions of euros); chain linked volume. Quarterly, calendar and seasonally adjusted. Based on the gross concept, i.e., both extra- and intra-area trade flows are accounted for. Source: AWM Database (code: *XTR*).

- *Investment (real).* Gross fixed capital formation (millions of euros); chain linked volume. Quarterly, calendar and seasonally adjusted. Source: AWM Database (code: *ITR*).

- *GDP (real).* Gross domestic product at market prices (million of euros); chain linked volume. Quarterly, calendar and seasonally adjusted. Source: AWM Database (code: *YER*).

- *Price level.* Harmonized index of consumer prices (HICP). Quarterly, adjusted for seasonal effects using the Census X-12 method (own calculations). Source: AWM Database (code: *HICP*).

- *Short-term nominal interest rate.* Euribor 3-month rate, last trade price; percent per annum. Source: AWM Database (code: *STN*). The Wu and Xia (2017) ECB shadow rate is used during the zero lower bound episode between 2008:Q4 and 2015:Q4.

B.4 Macro Data for Japan

Variables for Japan were retrieved from the FRED database of the Federal Reserve Bank of St. Louis and from the IMF IFS. The variables are

- *Business confidence.* The net percent of businesses that report increased confidence in their business performance related to present and future production of the company, order books (total and export), stocks of finished goods, tendency of future selling prices, and employment (see OECD, 2003). Quarterly, seasonally adjusted. Source: OECD Business Tendency Surveys for Manufacturing: confidence indicators: composite indicators: national indicator for Japan. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *JPNBSCICP02STSAQ*).

- *Effective exchange rate (real).* Trade-weighted real effective exchange rate (REER) index. An increase in REER implies that exports become more expensive and imports become cheaper; therefore, an increase indicates a loss in trade competitiveness. Details on the construction of the IMF's REER indices can be found in Bayoumi et al. (2005). Source: IMF IFS.

- *Exports (real).* Real exports of goods and services (billions of chained 2011 Yen). Quarterly, seasonally adjusted. Source: JP. Cabinet Office, National Accounts of Japan. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *JPNRGDPEGS*).
- *GDP (real).* Real gross domestic product (billions of chained 2011 Yen) Quarterly, seasonally adjusted. Source: JP. Cabinet Office, National Accounts of Japan, retrieved from FRED (code: *JPNRGDPEXP*).
- *Investment (real).* Gross fixed capital formation (chained 2010 Yen). Quarterly, seasonally adjusted. Source: OECD Quarterly National Accounts. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *JPNGFCFQDSNAQ*).
- *Price level.* Consumer price index: total, all items; index (2015=100). Quarterly, seasonally adjusted. Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *CPALTT01JPQ661S*).
- *Short-term nominal interest rate.* Yield on Japanese Treasury Bills; percent per annum. Quarterly average, not seasonally adjusted. Source: IMF IFS. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *INTGSTJPM193N*).

B.5 Macro Data for Mexico

Variables for Mexico were retrieved from the FRED database of the Federal Reserve Bank of St. Louis and from the IMF IFS. The variables are:

- *Business confidence.* The net percent of businesses that report increased confidence in their business performance related to present and future production of the company, order books (total and export), stocks of finished goods, tendency of future selling prices, and employment (see OECD, 2003); normalized (normal=100). Quarterly, seasonally adjusted. Source: OECD Business Tendency Surveys for Manufacturing: confidence indicators: composite indicators: OECD indicator for Mexico. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *BSCICP03MXM665S*).
- *Effective exchange rate (real).* Trade-weighted real effective exchange rate (REER) index. An increase in REER implies that exports become more expensive and imports become cheaper; therefore, an increase indicates a loss in trade competitiveness. Details on the construction of the IMF's REER indices can be found in Bayoumi et al. (2005). Source: IMF IFS.
- *Exports (real).* Exports of goods and services; index (2015=100). Quarterly, seasonally adjusted. Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *NAEXKP06MXQ661S*).
- *GDP (real).* Gross domestic product by expenditure in constant prices; index (2015=100). Quarterly, seasonally adjusted. Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *NAEXKP01MXQ661S*).

- *Investment (real).* Gross fixed capital formation; index (2015=100). Quarterly, seasonally adjusted. Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *NAEXKP04MXQ661S*).

- *Price level.* Consumer price index: total, all items; index (2015=100). Quarterly, adjusted for seasonal effects using the Census X-12 method (own calculations). Source: OECD Main Economic Indicators. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *MEXCPIALLQINMEI*).

- *Short-term nominal interest rate.* Yield on Mexican Treasury Bills; percent per annum. Quarterly average, not seasonally adjusted. Source: IMF IFS. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *INTGSTMXM193N*).

B.6 Macro Data for the United States

The U.S. variables were retrieved from the FRED database of the Federal Reserve Bank of St. Louis, the Bank for International Settlements (BIS), Wu and Xia (2016), and Gilchrist and Zakrajšek (2012). The variables are:

- *Capacity utilization: iron and steel.* Capacity utilization: durable manufacturing: iron and steel products; percent of capacity. Quarterly, seasonally adjusted. Source: Board of Governors of the Federal Reserve System. Retrieved from FRED, Federal Reserve Bank of St. Louis (code: *CAPUTLG3311A2SQ*).

- *Commodity prices.* The weighted sum of oil prices and non-oil commodity prices in U.S. dollars. Source: AWM Database (code: *COMPR*).

- *Effective exchange rate (real).* Real effective exchange rate (narrow) index. Trade weighted average of a basket of bilateral exchange rates, adjusted with the corresponding relative consumer prices. Quarterly average, not seasonally adjusted. An increase indicates an appreciation. Source: BIS (code: *RNUS*).

- *GDP (real).* Real gross domestic product in billions of chained 2009 Dollars. Quarterly, seasonally adjusted annual rate. Source: U.S. Bureau of Economic Analysis. Retrieved from FRED (code: *GDPC1*).

- *GZ credit spread.* Corporate bond spread index constructed by Gilchrist and Zakrajšek (2012). Quarterly average, percent, not seasonally adjusted. Retrieved from the web page of Simon Gilchrist (http://people.bu.edu/sgilchri/).

- *Hours.* Average weekly hours per worker of production and nonsupervisory employees: manufacturing. Quarterly, seasonally adjusted. Source: U.S. Bureau of Labor Statistics. Retrieved from FRED (code: *AWHMAN*).

- *Imports (real).* Real imports of goods and services, billions of chained 2012 Dollars. Quarterly, seasonally adjusted annual rate. Source: U.S. Bureau of Economic Analysis. Retrieved from FRED (code: *IMPGSC1*).

- *Sales (real).* Real manufacturing and trade industries sales. Millions of chained 2012 Dollars. Quarterly, seasonally adjusted. Source: Federal Reserve Bank of St. Louis. Retrieved from FRED (code: *CMRMTSPL*).

- *Short-term nominal interest rate.* Effective federal funds rate, in percent. Quarterly average, not seasonally adjusted. Source: Board of Governors of the Federal Reserve System. Retrieved from FRED (code: *FEDFUNDS*). To account for the stance of monetary policy during the zero lower bound episode, the Wu and Xia (2016) shadow federal funds rate is used between 2009:Q1 and 2015:Q4.

- *Treasury bond rate (real).* The real Treasury bond rate is calculated following Basset et al. (2014) as the nominal 10-year Treasury constant maturity rate minus one-year-ahead expectations of CPI inflation, as reported by the Survey of Professional Forecasters of the Federal Reserve Bank of Philadelphia. The quarterly 10-year Treasury constant maturity rate (in percent, not seasonally adjusted) comes from the Board of Governors of the Federal Reserve System, retrieved from FRED (code: *GS10*).

- *Unemployment rate.* Civilian unemployment rate: number of unemployed as a percentage of the labor force. Quarterly, seasonally adjusted. Source: U.S. Bureau of Labor Statistics, retrieved from FRED (code: *UNRATE*).

- *Wages (real).* Nominal wages deflated by the CPI. Nominal wages: indexes of aggregate weekly payrolls of production and nonsupervisory employees: manufacturing, index 2002=100. Quarterly, seasonally adjusted. Source: U.S. Bureau of Labor Statistics. Retrieved from FRED (code: *CES3000000035*). CPI: consumer price index for all urban consumers: all items, index 1982-1984=100. Quarterly, seasonally adjusted. Source U.S. Bureau of Labor Statistics. Retrieved from FRED (code: *CPIAUCSL*).

C Robustness Checks

This section presents the results of robustness checks concerning the first-stage regression and the euro area VAR. Robustness checks on the VARs for other countries are available upon request from the author.

C.1 Robustness to the Presence of Measurement Error in the Shock Series

Since exogenous shocks to U.S. trade policy are not directly observed, I derive proxies from data on U.S. TTB policies. The construction of such proxies is subject to measurement error which might ultimately bias the VAR estimates. Identification in the benchmark VAR is carried out using the Cholesky decomposition in a recursive VAR with the shock measure ordered first. If the shock series is contaminated with measurement error that is independent from the underlying structural shock, Plagborg-Møller and Wolf (2019) show that ordering the shock measure first in a recursive VAR provides valid structural estimation. Hence, even if the proxies derived from data on U.S. TTB policies are noisy measures of the true underlying shocks, valid estimates of the impact of TTB protection can be obtained by including the externally identified shock series into the VAR, as long as the shock is ordered first.

Nevertheless, I also employ two alternative methods to estimate impulse responses. First, I use the proxy-VAR approach proposed by Stock and Watson (2012) and Mertens and Ravn (2013). This treats the potentially noisy shock measure as an external instrument for the underlying trade policy shock. The IRFs are then recovered from a VAR model for EA variables using instrumental variable methods (see also Stock and Watson, 2018). IRFs from this method are identified up to scale. Therefore, the IRFs are normalized with respect to the impact effect on business confidence. The impulse responses estimated using the external instruments approach are statistically indistinguishable from the benchmark IRFs at the 90% level, suggesting that the contractionary effects of trade policy announcements do not hinge on the recursive identification scheme (see Figure A.1, top panel).

Second, I estimate impulse responses by the local projection method proposed by Jordà (2005), using the baseline indicator directly as a shock measure. Ramey (2016) argues that local projections are useful as a heuristic check on VAR models. The following regression is estimated:

$$z_{t+h} = \alpha + \theta_h \xi_t + B(L) Z_t + \eta_{t+h},$$

where z_t are endogenous euro area variables ($z_t \in Z_t$), θ_h is the estimate of the impulse response of z_t at horizon h to shock ξ_t, where ξ_t is the baseline indicator of cyclically adjusted U.S. import protection. $B(L) = I + \beta_1 L + ... + \beta_p L^p$ is a polynomial in the lag operator, and α is a constant term. The regression is estimated with OLS in (log-)levels, and standard errors are corrected for heteroscedasticity and serial correlation using the Newey and West (1987) estimator. Local projections provide an arguably more flexible approach to the estimation of IRFs than VARs because they impose fewer restrictions than standard methods. This comes at the cost that the IRFs are less precisely estimated and are somewhat more erratic, as illustrated in Figure A.1 (bottom panel). Yet, the main results obtained using the recursive VAR carry over to the local projection framework.

Consider now the case in which the shock series is contaminated by measurement error that is related to the underlying structural shock. For instance, this would be the case if TTB investigations had been initiated at the same time as other types of trade protection measures. The systematic coincidence of TTBs with other policy actions might bias the VAR estimates. Using data from the United Nations Trade Analysis Information System database, I add to the set of control variables in the first-stage regression a variable capturing the quarterly count of U.S. non-tariff barrier (NTB) measures initiated during the sample period (Non-Tariff-Barrier Count). NTBs include technical barriers to trade, export-related measures, price and quantity

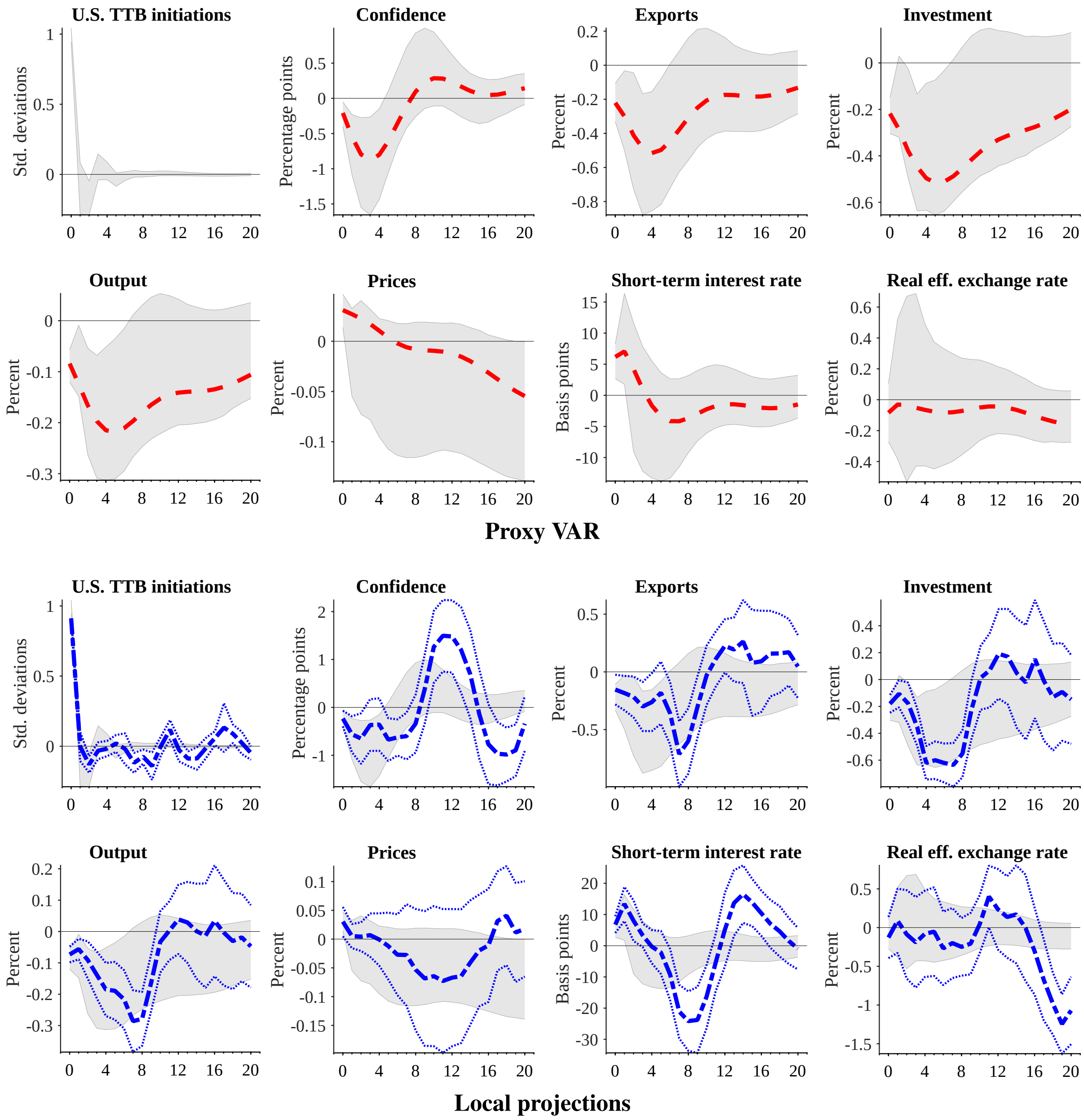


Figure A.1: **Effects of a U.S. Trade Policy Announcement Shock: Proxy-VAR and Local Projection Estimates**

Notes: Top panel: Proxy-VAR estimates of median impulse responses (red dashed lines) of euro area variables to a one-standard-deviation U.S. trade policy announcement shock. U.S. TTB initiations are measured by the real trade value of TTB initiations adjusted for macroeconomic factors. Bottom panel: Local projection estimates of impulse responses to an orthogonalized one-standard-deviation increase in the real trade value of U.S. TTB initiations adjusted for macroeconomic factors (blue dashed-dotted lines), with 90% confidence intervals based on Newey and West (1987) heteroscedasticity and autocorrelation consistent standard errors (blue dotted lines). Impulse responses are estimated for 20 quarters. The figures also depict the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

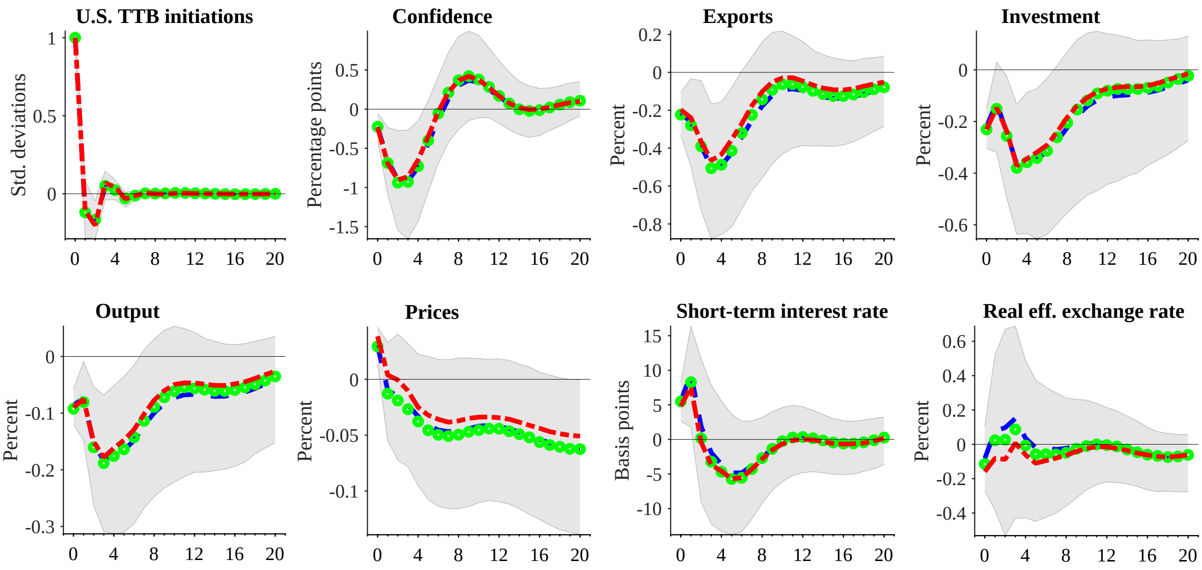


Figure A.2: **Effects of a U.S. Trade Policy Announcement Shock: Robustness**

Notes: Median impulse responses of euro area variables to a one-standard-deviation U.S. trade policy announcement shock, estimated for 20 quarters using: (1.) the residual from an OLS regression that includes, in addition to the baseline specification, a variable measuring the number of non-tariff-barriers initiated by the U.S. in each quarter (green circles); (2.) the residual from an OLS regression that includes, in addition to the baseline specification, a dummy variable that takes the value of one in years of U.S. presidential elections (blue dashed lines); and (3.) the residual from an OLS regression that includes, in addition to the baseline controls, a variable measuring the co-movement between the annual growth rates of U.S. and RoW GDP (red dashed-dotted lines). The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

control measures, sanitary and phytosanitary measures and preshipment inspection. Purging the shock series from this variable removes systematic measurement error that might be related to the shock of interest. I find a weakly negative and statistically significant correlation between the initiation of TTB and NTB measures (see column (1) of Table A.2). However, using the residual shock series that is cleansed of NTBs yields VAR estimates that are nearly identical to the baseline results (see Figure A.2). Hence, the VAR estimates are robust to measurement error in the shock series that might arise from the coincidence with other trade policies.

C.2 Robustness to the U.S. Electoral Cycle

There is a long-standing literature analyzing the relation between politics and trade policy. For instance, Finger et al. (1982) study the role of political factors in the administrative mechanisms associated with TTB protection, while Grossman and Helpman (1994) model situations in which special-interest groups attempt to influence an incumbent government's choice of trade policy. The Financial Times article cited in the main text on the 2001 steel safeguards argues that "*President George W. Bush's [...] threat to curb steel imports and demand for international talks to cutting overcapacity and subsidies are a cave-in to special interest lobbying and*

	(1) NTBs (OLS)	(2) Election (OLS)	(3) Co-move (OLS)	(4) Imposed (OLS)	(5) Anti-dump (OLS)	(6) Ongoing (OLS)	(7) Initiated (OLS)
Lagged Dependent	-0.152**	-0.145**	-0.177**	-0.101**	-0.152**	-0.142**	-0.143**
	(0.030)	(0.034)	(0.035)	(0.038)	(0.054)	(0.036)	(0.024)
ΔGDP	0.251	0.171	0.556	-0.255	0.713	0.244	0.648
	(0.397)	(0.291)	(0.324)	(0.170)	(0.414)	(0.482)	(0.361)
ΔSales	0.013	0.022	-0.145	0.267*	-0.276	-0.057	0.305
	(0.243)	(0.203)	(0.263)	(0.105)	(0.251)	(0.325)	(0.272)
ΔWage	-0.229	-0.279	-0.242	-0.316	-0.092	-0.417	-0.435*
	(0.231)	(0.180)	(0.200)	(0.228)	(0.188)	(0.331)	(0.213)
ΔUnemployment	-1.088	-1.292	-0.844	-1.193	-0.883	-2.097	-1.077
	(1.143)	(0.963)	(0.907)	(0.994)	(0.942)	(1.471)	(0.715)
ΔHours	-0.084	-0.074	-0.133	-0.292	0.032	-0.314	-0.522
	(0.305)	(0.308)	(0.275)	(0.173)	(0.269)	(0.307)	(0.423)
Capacity Utilization (Steel)	-0.046	-0.054	-0.046	-0.057	-0.070*	-0.106	-0.141
	(0.040)	(0.034)	(0.034)	(0.037)	(0.034)	(0.054)	(0.079)
Real 10Y T-Bond	0.395	0.441	0.243	0.393*	0.288	0.363	0.587*
	(0.311)	(0.279)	(0.213)	(0.190)	(0.301)	(0.320)	(0.232)
GZ Credit Spread	0.415	0.389	0.039	0.539*	0.038	0.480	0.588
	(0.487)	(0.533)	(0.413)	(0.306)	(0.379)	(0.577)	(0.344)
Global Activity Index	0.001	0.001	-0.078	0.008	0.004	0.002	-0.003
	(0.005)	(0.004)	(0.045)	(0.004)	(0.004)	(0.006)	(0.006)
ΔREER	-0.042	-0.040	0.034	-0.044	-0.053	-0.016	-0.085*
	(0.054)	(0.054)	(0.081)	(0.061)	(0.047)	(0.069)	(0.042)
ΔImports	-0.003	0.018	-0.025	0.081	-0.034	0.071	0.057
	(0.067)	(0.057)	(0.017)	(0.082)	(0.053)	(0.106)	(0.086)
ΔCommodity Prices	-0.013	-0.014	-0.002	-0.023	-0.009	-0.013	-0.009
	(0.019)	(0.017)	(0.005)	(0.012)	(0.015)	(0.023)	(0.020)
Non-Tariff-Barrier Count	-0.004**						
	(0.001)						
Election Year Dummy		0.195					
		(0.679)					
ΔGDP Co-movement			0.786**				
			(0.216)				
R-squared	0.072	0.066	0.105	0.104	0.083	0.100	0.109

Table A.2: **Regression Estimates of U.S. TTB Protection on Macro Factors - Robustness**

Notes: Column (1): OLS estimates with RHS: baseline specification and count of U.S. non-tariff barrier measures. Column (2): OLS estimates with RHS: baseline specification and election year dummy. Column (3): OLS estimates with RHS: baseline specification and GDP growth co-movement between U.S. and RoW. Column (4): OLS estimates with LHS: trade value indicator with product count derived exclusively from cases that led to the final imposition of TTBs. Column (5): OLS estimates with LHS: trade value indicator with product count derived exclusively from anti-dumping cases. Column (6): OLS estimates with LHS: trade value indicator with product count derived by including also products for which a TTB is ongoing or already in place at the time of the investigation. Column (7): OLS estimates with LHS: trade value indicator with product count derived by including all cases that were initiated, regardless of subsequent withdrawal or termination. HAC standard errors are in parentheses. Model includes a constant term whose estimate is suppressed. Asterisks ** and * indicate statistical significance at the 1 and 5 percent levels, respectively. Sample: 1988:Q2-2015:Q4.

an admission that he has lost control of the trade agenda to protectionist forces in Congress. [...]" – See: "Bush's lack of steel" Editorial in: Financial Times (London, England), Issue 34,546, p. 20, Thursday, June 7, 2001; retrieved from the Financial Times Historical Archive, 1888-2010 (Gale). The question thus arises whether the political cycle plays a role for the use of TTBs. I include into the first-stage regression a dummy variable that takes the value of one during U.S. presidential election years (Election Year Dummy). However, I do not find statistically significant evidence that the initiation of TTBs is more common right before U.S. presidential elections (see column (2) of Table A.2). Moreover, using the residual shock series that is purged of the U.S. electoral cycle does not lead to any material changes in the VAR estimates (see Figure A.2).

C.3 Robustness to U.S. and Global Business Cycle Co-movement

Is it possible that what the VAR is picking up is the synchronicity of the U.S. and global business cycles that are just not being captured by Equation (1)? I augment the set of controls with a measure of business cycle co-movement that captures synchronicity between the U.S. and the global business cycle, in order to verify whether controlling for the latter plays a material role for the VAR estimates. Following Kalemli-Ozcan et al. (2013), this variable is defined as the negative of the absolute value of differences between the growth rates of U.S. real GDP, $Y_t^{U.S.}$, and a composite variable representing real GDP for the "rest of the world" (ROW), Y_t^{ROW}, computed as: Δ GDP Co-movement$_t$ = $-|(\log Y_t^{U.S.} - \log Y_{t-4}^{U.S.}) - (\log Y_t^{ROW} - \log Y_{t-4}^{ROW})|$. Values closer to zero indicate higher synchronicity. Following Enders and Mueller (2009), GDP growth for the ROW is defined as the weighted average of annual GDP growth rates for the euro area, the U.K., Japan, and Canada. The weights are calculated at annual purchasing power parity (PPP) values, based on data from the Penn World Tables. Indeed, I find that more synchronous U.S. and ROW business cycles are associated with more TTB protection (see column (3) of Table A.2). However, the VAR estimates obtained by using the residual shock series adjusted for business cycle co-movement are virtually identical to the baseline results (see Figure A.2).

C.4 Robustness to the Assumptions Regarding Aggregation of TTB Cases

The assumptions made during the aggregation of micro data on TTB cases into composite time series measures closely follows Bown and Crowley (2013). I conduct four robustness checks to test for sensitivity of the baseline results with respect to these assumptions. First, I compute the trade value of TTB initiations by excluding cases in which injury or unfair trade allegations were proven unfounded, as well as cases that were withdrawn by the petitioner or terminated by the government agency prior to ruling, leaving only those investigations that conclude with

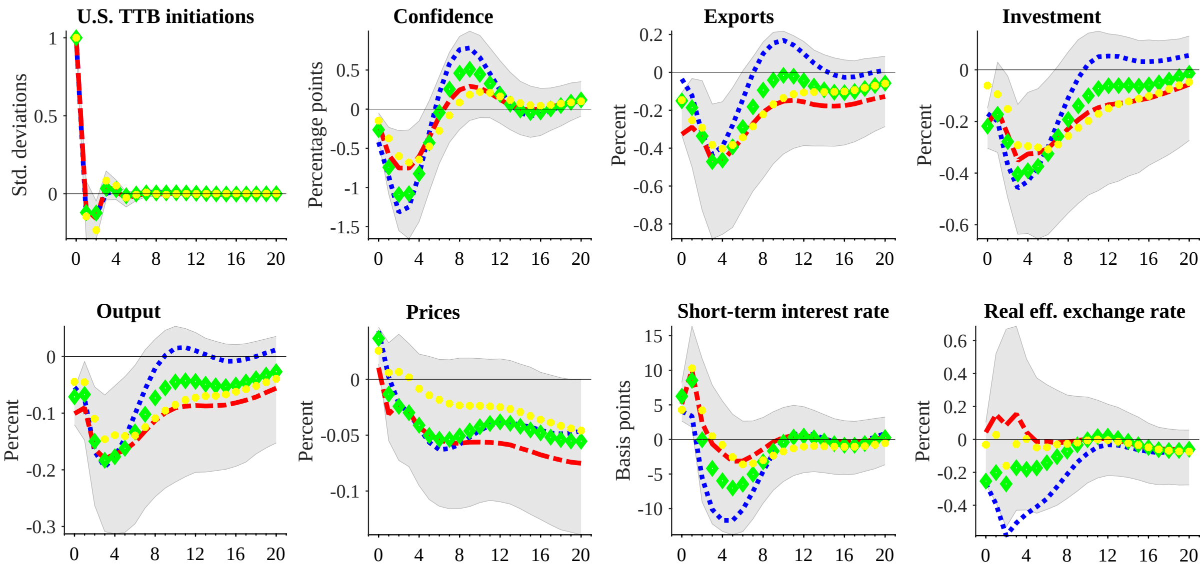


Figure A.3: **Effects of a U.S. Trade Policy Announcement Shock: Robustness**

Notes: Median impulse responses of euro area variables to a one-standard-deviation U.S. trade policy announcement shock estimated for 20 quarters, using as shock measure: (1.) the macro-adjusted trade value based on the count of products subject to new TTB investigations that led to an affirmative final decision (blue dotted lines); (2.) the macro-adjusted trade value based on the count of products subject to new anti-dumping investigations (red dashed-dotted lines); (3.) the macro-adjusted trade value based on the count of products subject to new TTB investigations, including cases in which new trade protection initiatives fall into the same HS06 category as a previously imposed TTB or that were subject to a simultaneous or previously imposed barrier under a different TTB measure (green diamonds); and (4.) the macro-adjusted trade value based on the count of products subject to new TTB investigations, including cases that were terminated prior to a preliminary injury determination or due to the latter being negative (yellow stars). The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

formal trade barriers being imposed (see specification (4) in Table A.2 for OLS estimates of a regression of this variable on the baseline controls). Second, I derive the trade-value indicator exclusively from anti-dumping cases (see specification (5) in Table A.2). Third, the trade-value indicator is derived by including cases regardless of whether another investigation on the same products is ongoing or a TTB is already in place at the time of initiation (see specification (6) in Table A.2). Finally, the trade-value indicator is constructed on the broadest possible basis by including all cases that were initiated, regardless of subsequent withdrawal or termination at any stage of the investigation (see specification (7) in Table A.2).

The regression results in Table A.2 show evidence of first order autocorrelation and some signs of a statistically significant link between import protection and macroeconomic fluctuations that are in line with the baseline estimates. In particular, I find a significant negative relationship between capacity utilization in the steel industry and anti-dumping activity, as well as between TTB initiations and wage growth. I also find some evidence that tightening external financing constraints and a real depreciation of the U.S. dollar are significantly associated with more TTB protection. Figure A.3 shows that the differences between using the baseline

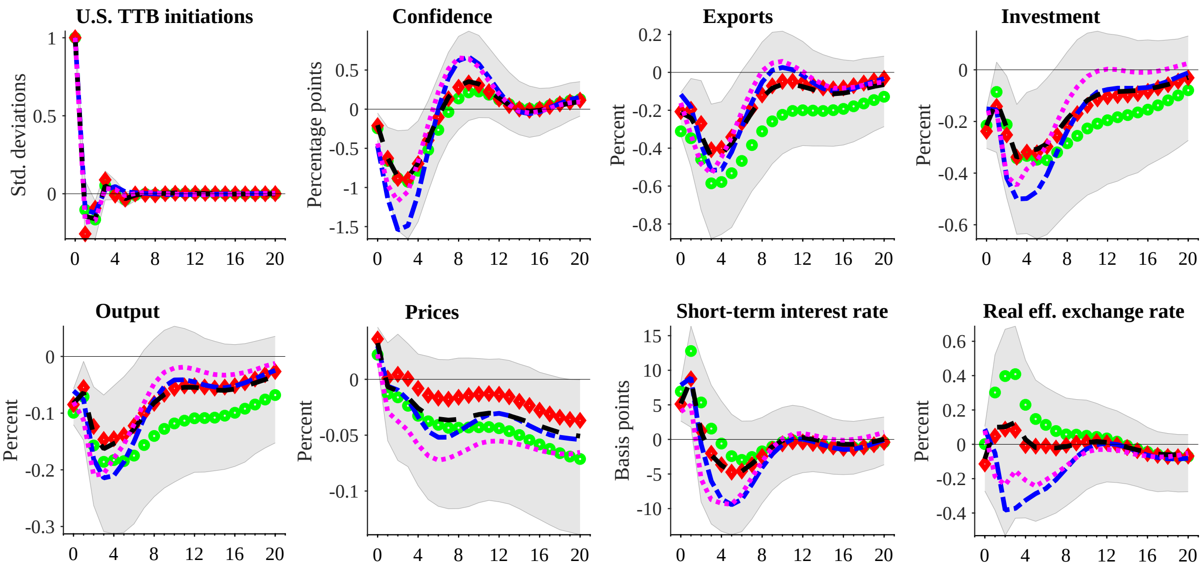


Figure A.4: **Effects of a U.S. Trade Policy Announcement Shock: Robustness**

Notes: Median impulse responses of euro area variables to a one-standard-deviation U.S. trade policy announcement shock estimated for 20 quarters, using as shock measure: (1.) the residual obtained from a Tobit model instead of an OLS regression (black dashed lines); (2.) the baseline residual expressed in relation to aggregate U.S. imports instead of in real terms (green circles); (3.) the real trade value without adjustment for macro factors (red diamonds); (4.) the unweighted count of products subject to new TTB initiations, i.e., the Bown-Crowley measure of time-varying import protection (blue dashed-dotted lines); and (5.) the Bown-Crowley measure purged of macro factors (magenta dotted lines). The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

shock series corresponding to the standardized residual from specification (2) in Table 4 and alternative residuals derived from specifications (4)-(7) in Table A.2 are quantitatively small and overwhelmingly not statistically significant.

C.5 Robustness to Some Additional Alternative Shock Series

I replace the baseline shock series with five additional alternative shock series. These are: (1.) the residual obtained from Tobit estimates of the baseline regression specification; (2.) the residual from the baseline regression with trade values expressed in relation to aggregate U.S. imports instead of the GDP-deflator; (3.) the real trade value without adjustment for macro factors; (4.) the count of products subject to new TTB initiations, i.e., the Bown-Crowley measure of time-varying import protection; and (5.) the Bown-Crowley measure purged of macro factors. Using these five alternative shock series produces VAR estimates that are qualitatively identical and quantitatively close to the baseline results (see Figure A.4).

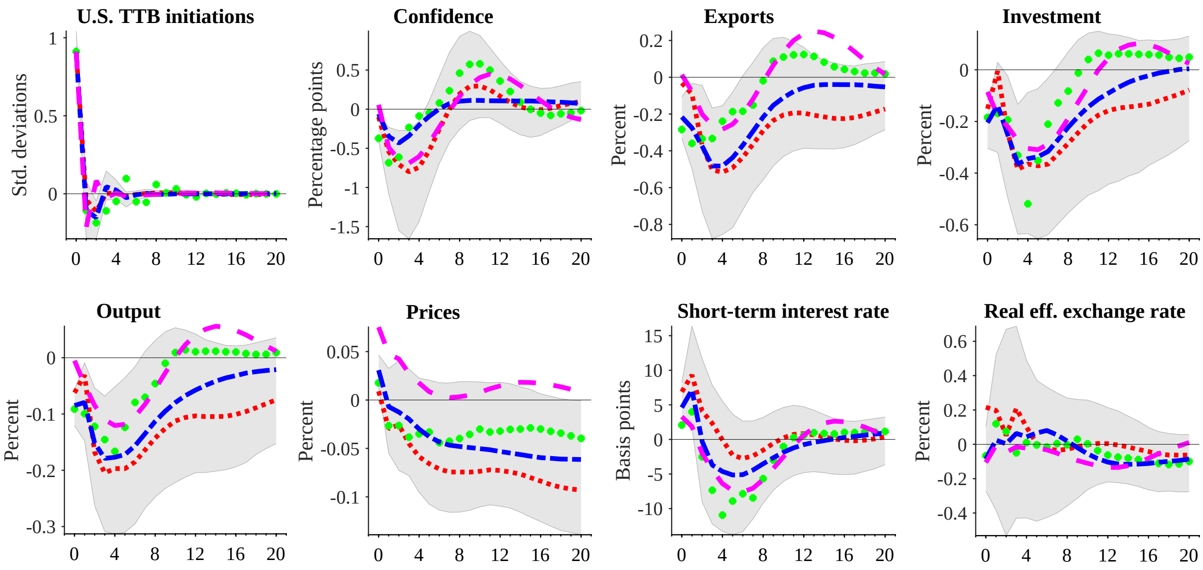


Figure A.5: **Effects of a U.S. Trade Policy Announcement Shock: Robustness**

Notes: Median impulse responses of euro area variables to a one-standard-deviation U.S. trade policy announcement shock estimated for 20 quarters. The baseline VAR is modified as follows: (1.) the two largest shock events (1989:Q3 and 2001:Q2) are dropped from the sample (red dotted lines); (2.) business confidence is replaced by a survey-based measure of consumer confidence (blue dashed-dotted lines); (3.) the benchmark VAR estimated between 1999:Q1-2015:Q4 with $p = 1$ (magenta dotted lines); and (4.) the reduced-form VAR is estimated with four lags (green crosses). The figure also depicts the 90% confidence intervals from the benchmark VAR (grey areas); see also notes to Figure 4. Sample: 1988:Q2-2015:Q4.

C.6 Robustness to Large Shocks, Consumer Confidence, Sample Period, and Lag Structure

I carry out four further robustness checks shown in Figure A.5. First, I verify whether the two largest shocks (in 1989:Q3 and 2001:Q2) are the sole drivers of the empirical findings. To that aim, I omit them from the baseline measure by setting the value of the shock series in those quarters to zero. The omission of these two shock events does not have a material impact on the estimated impulse responses. Second, I replace the survey-based measure of business confidence with a comparable measure of consumer confidence. The correlation coefficient between the two measures is relatively high at 0.67. The macro effects are robust to adding consumer confidence to the VAR. The reaction of consumer confidence to a trade policy announcement shock is somewhat weaker than the response of business confidence; yet, the difference is statistically insignificant. Hence, the two measures seem to capture the same underlying expectations mechanism. Third, I estimate the benchmark VAR for the post-1999 era because the euro area did not exist for the first ten years of the sample. Due to the relatively small sample size, the VAR is estimated with one lag of the endogenous variables. VAR estimates obtained for the period between 1999-2015 are overwhelmingly statistically indistinguishable from the benchmark estimates at the 90% level. Dropping the pre-1999 period from the sample does not

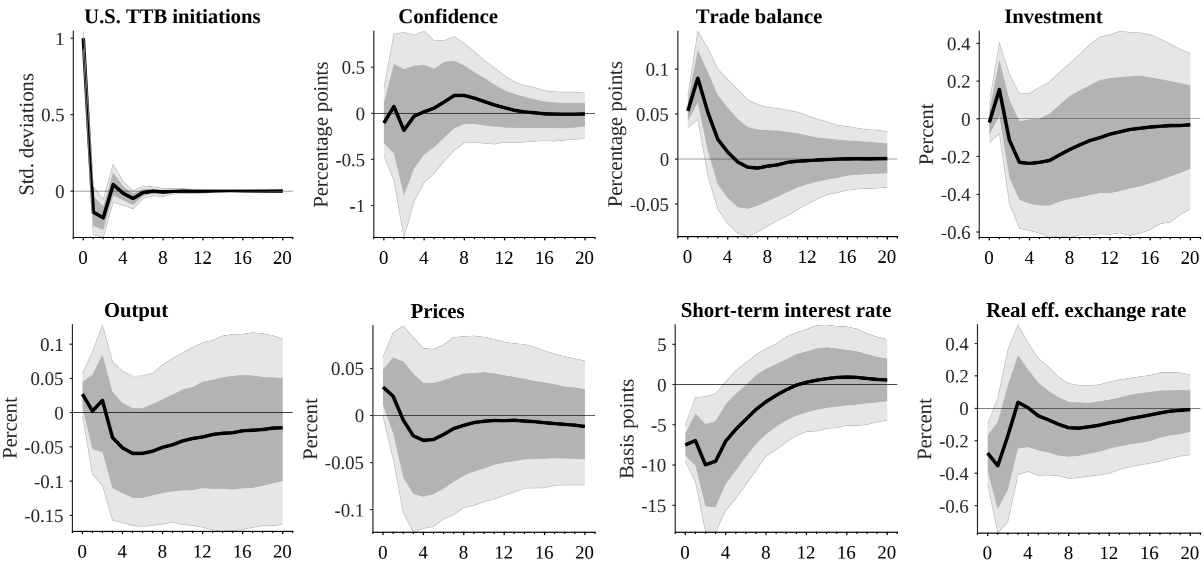


Figure A.6: **Effects of a U.S. Trade Policy Announcement Shock: United States**

Notes: Estimated impact of a one-standard-deviation U.S. trade policy announcement shock on the United States. Impulse responses are estimated for 20 quarters. U.S. TTB initiations are measured by the real trade value of TTB initiations adjusted for macroeconomic factors. Shaded areas represent the 68% (dark grey) and 90% (light grey) confidence intervals based on 2,000 bootstrap replications. Sample: 1988:Q2-2015:Q4.

materially alter the main conclusions. Finally, I show that the results are robust to estimating the benchmark VAR model with four lags in the reduced-form system.

C.7 VAR Estimates for the United States

Finally, Figure A.6 shows the estimated impact of a one-standard-deviation U.S. trade policy announcement shock on U.S. variables. Importantly, I find that the shock does not provide a significant stimulus to the U.S. economy. Instead output, investment, and confidence hardly display any significant reaction to a surprise increase in U.S. trade protection. The U.S. trade balance significantly improves, in line with the decline in foreign exports estimated for the euro area, China, Japan, and Mexico. The U.S. short term interest rate decreases and the REER depreciates.