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Hidden Protectionism? Evidence from Non-Tariff barriers to trade in the United States

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Hidden Protectionism? Evidence from Non-tariff Barriers to Trade in the United States

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Abstract

Can the enforcement of product standards be protectionism in disguise? This paper estimates the costs of non-compliance with U.S. product standards, using a new database on U.S. import refusals from 2002 to 2014. We find that import refusals decrease exports to the United States. This trade reducing effect is driven by developing countries and by refusals without any product sample analysis, in particular during the Subprime Crisis and its aftermath. We also provide evidence that given product standards have been enforced more strictly during the crisis. These results are consistent with the existence of counter-cyclical, hidden protectionism due to non-tariff barriers to trade in the United States.

JEL: F13, F14, O24, F63.

Keywords: Hidden protectionism, international trade, developing countries, import refusals, regulatory costs, disaggregated, United States.

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1. Introduction

Non-tariff barriers to trade (NTBs) like product standards and technical regulations have increased in importance compared to tariffs that are at historical lows.[1] The discussion on NTBs is characterized by two opposing trends. While some countries aim at harmonizing product standards to reap further gains from trade (e.g., trade talks between the United States and European Union), fear of protectionism has led to a close monitoring of NTBs worldwide during the Great Recession (e.g., Baldwin and Evenett, 2009).

Product standards are imposed to overcome market failures and protect the health of domestic consumers. In the United States, the Food and Drug Administration (FDA) is responsible for ensuring the safety of domestic and foreign products, i.e., to enforce U.S. product standards. Those products not complying with U.S. product standards are refused entry into the market by the FDA. For domestic products, the FDA needs to prove a violation to deny access to the U.S. market. In the case of imports, however, the FDA has the authority to refuse import shipments that "[only] appear to violate" a certain U.S. product standard (Liu 2010). This formulation in Section 801(a) of the Federal Food, Drug, and Cosmetic (FD&C) Act leaves considerable room for discriminatory action of FDA officials with respect to imports. We analyse this important issue by collecting a new data set that combines disaggregated import data from the U.S. International Trade Commission (ITC) with import refusals from the FDA.

In this paper, we estimate the impact of U.S. product standards enforcement on trade at different times of the business cycle. Our main variable of interest, import refusals, is a *de facto measure for technical regulations*. This variable captures all incidences of non-compliance with U.S. product standards, i.e., when a given legal technical regulation indeed bites. We show that the negative impact of import refusals on imports to the United States is substantial for poorer countries. While OECD countries are largely unaffected, our estimates imply that a one standard deviation increase in refusals reduces short- and long-run exports from an average non-OECD country by USD 2.8 to 5 billion. By examining the type of inspection that underlies a given import refusal, we show that the trade reducing effect is mainly triggered by refusals without any product sample analysis during the Subprime Crisis

[1] Similarly, Baldwin's famous quote says that "[t]he lowering of tariffs has, in effect, been like draining a swamp. The lower water level has revealed all the snags and stumps of non-tariff barriers that still have to be cleared away." (Baldwin, 1970, quoted in Baldwin, 2000).

and its aftermath. Furthermore, we use U.S. notifications to the WTO SPS and TBT agreements to control for changes in U.S. product standards regulation. The results show that the existing U.S. product standards have been enforced more strictly during the crisis. We conclude that these results for the United States are consistent with (but do not prove) hidden protectionism due to *stricter enforcement of product standards*.

We estimate a bilateral gravity model for 93 imported product-groups to the United States for the years 2002 to 2014. We proceed in three steps by reporting OLS estimates, standard fixed effect estimates and, then, dynamic panel estimates. The last and preferred specification does not only allow us to control for past import flows and use lagged import refusals as internal instruments for our non-compliance measure, but we can also extend the framework to include additional, external instruments drawn from the EU - Rapid Alert System for Food and Feed (RASFF) database. Since the notifications collected by the RASFF from participating countries are plausibly exogenous to U.S. import demand, but likely to be correlated with U.S. refusals, they constitute a valid instrument.

Why are thousands of shipments blocked from entering the U.S. market each year? The FDA names two main reasons for import refusals: adulteration and misbranding. Recent reports on blocked U.S. imports of toys containing lead fall in the first category with products being inferior and entailing substantial health risks.[2] But adulteration can also simply stem from differing product standards between trading countries. Second, a product might be denied entry into the United States due to misbranding, i.e., U.S. labelling standards are not met or necessary certificates for conformity assessment are not provided by the exporter.

The FDA might be also subject to lobbying and political pressures.[3] In one of its most controversial moves, the FDA issued an outright ban of all grapes from Chile in March 1989 due to a non-lethal contamination of two grapes with cyanide (Engel, 1999; Hawthorne, 2005). It remains unclear to date, whether the FDA simply overreacted or the U.S. government aimed for a weakening of the Pinochet regime. Similar to technical regulations (see Trefler, 1993 and Essaji, 2008), stricter enforcement of product standards may be

[2] Decker, Brett and William Triplett "China's Poisonous Exports: PRC Products Aren't just Cheap, They're Dangerous," The Washington Times, November 16th, 2011.

[3] After all, the FDA is a government agency, its commissioner a political appointee and the revolving door also spins at the FDA. For the U.S. Department of Defense, Luechinger and Moser (2014) show that conflicts of interest can arise due to the revolving door. The Government Accountability Office indeed acknowledges a staff turnover rate at the FDA above the federal government average in 2002 (cited in Hawthorne, 2005, p. 30).

imposed due to protectionist motives in the United States. Lamb (2006) provides anecdotal evidence that political pressure from U.S. avocado producer associations has been driving the boycott of Mexican Hass avocados until 1997. More recently, U.S. catfish producers have lobbied for more frequent inspections of catfish imports to protect their industry.[4,5] In 2015, Canada and Mexico were authorized by the WTO to retaliate against the meat labelling requirement (COOL) that the U.S. had put into force at the height of the financial crisis in 2009.[6] According to Watson and James (2013) regulatory protectionism exists in the United States and Baldwin (2000) is especially concerned about its effect on developing countries. In a similar vein, Staiger (2015) states that developed countries' behind-the-border NTBs like technical regulations represent an important NTB for developing countries' exporters.

Graph 1 sheds some first light on stricter enforcement of U.S. product standards during the Subprime Crisis and its aftermath. This figure shows that the total number of shipments inspected by the FDA increased sharply from less than 140,000 in 2008 to close to 280,000 in 2011, staying close to its peak in the years thereafter. These FDA-inspections include inspections with and without a product sample analysis. Strikingly, after a strong rise, incidences of non-compliance with U.S. product standards fell again hand in hand with the decreasing unemployment rate. This is even more striking if one considers that imports have resurged after the crisis (Graph 2) and that the product quality of U.S. imports has not changed during the crisis and its aftermath (Levshenko et al. 2011). A further remarkable fact is that import refusals without any product sample analysis dominate the evolution of total import refusals. Our regression analyses will elaborate on this type of inspection that is arguably prone to potential hidden protectionism.

Our paper contributes to the existing literature in several ways. First, we contribute to the recent empirical literature on protectionism (see Rose, 2013; Bown and Crowley, 2013; Kee et al., 2013) by highlighting another channel through which governments might temporarily seek import protection: a stricter enforcement of product standards. While *de jure* changes in product standards need to be notified to the WTO, *de facto* changes in given product standards such as stricter border inspections might fly below the radar of

[4] Nixon, Ron, "Number of Catfish Inspectors Drives a Debate on Spending," The New York Times, July 26th, 2013.
[5] Jouanjean (2012) provides evidence that U.S. producer associations influence U.S. market access regulation for imports of fresh fruits and vegetables. As exemplified for Russia, product standards and stricter inspections at the border might be even used for foreign policy purposes. See Kramer, Andrew, "Chocolate Factory, Trade War Victim," The New York Times, October 29th, 2013; Herszenhorn, David, "Russia Putting a Strong Arm on Neighbors," The New York Times, October 22th, 2013.
[6] Miles and Nickel "Canada, Mexico win $1 billion sanctions against U.S. in meat spat", Reuters, Dec 7th, 2015.

international policy makers. Note that the U.S. jurisdiction allows an imported product to be refused entry "if it [only] appears to violate" a certain U.S. product standard (Section 801(a) of the Federal Food, Drug, and Cosmetic (FD&C) Act). We are to the best of our knowledge the first to consider the type of inspection leading to import refusals and to link the effect of import refusals to the business cycle. Second, we add to the trade and development literature by quantifying the short- and long-run trade costs due to non-compliance with U.S. product standards. Most alternative measures of product standards, e.g., notifications to the WTO, are based on technical regulations that are most favoured nation (MFN) measures without variation across exporters. In contrast, our measure substantially varies across countries, product-groups and time. This allows us to factor in that any potential import protection is trading partner- and product-specific and to control for country-product-specific factors that are often omitted in other studies. Third, we contribute to the literature on import refusals by demonstrating how important it is to account for the endogenous nature of refusals.[7] Thereby, endogeneity can arise due to import protection or risk-guided inspections. Product-groups with increasing imports are more closely monitored and inspected by the FDA.

The remainder of the paper is organized as follows. Sections 2 and 3 discuss the related literature, the institutional background of the U.S. Food and Drug Administration (FDA) and some descriptive statistics for the new import refusal database. Section 4 provides a description of our data set and an outline of the empirical strategy. Section 5 presents the empirical results and Section 6 offers concluding remarks.

2. Related Literature

Our paper is related to the literature on endogenous protectionism, counter-cyclical protectionism, the theory on product standards and the effect of non-tariff barriers to trade, in particular for developing countries.

In a seminal paper, Trefler (1993) argues that the level of trade protection is not exogenous but increasing in import competition and domestic lobbying efforts. Similarly, the paper closest to our study is Essaji (2008) who analyzes the effects of technical regulations

[7] In contrast to other studies, we quantify the impact of import refusals not only for food products, but also for pharmaceuticals, cosmetics and manufacturing products. This is a smaller contribution to the literature.

for a cross-section of sectoral trade flows to the United States in 1999. Note that Essaji (2008) uses a *de jure* measure that captures the existence of technical regulations in a given sector, but does not seek to distinguish between trade increasing and reducing regulations. To address potential endogeneity of technical measures, Essaji (2008) instruments U.S. technical regulations with such regulations of countries with similar regulatory processes, but different import patterns. Both important contributions clearly show that the effects of protectionism are economically large, once endogeneity is taken into account.

Bagwell and Staiger (2003) provide a theoretical model, where trade policy becomes more protectionist in bad economic times. The more recent empirical evidence on this issue is mixed. While Rose (2013) argues forcefully that protectionism – as measured by a broad set of tariff and non-tariff barriers – has neither been counter-cyclical in the United States nor worldwide after World War II, Bown and Crowley (2013) and Kee et al. (2013) offer a more nuanced picture. Kee et al. (2013) conclude that only a few countries have markedly increased their tariffs from 2008 to 2009, but the relatively modest U.S. trade policy reaction has been an NTB, namely antidumping.[8] Bown and Crowley (2013) investigate the relationship between the business cycle and another NTB for five OECD countries. The most relevant result for our paper: Bown and Crowley (2013) provide evidence that the number of disaggregated product groups affected by temporary trade barriers increases with negative macroeconomic shocks in the United States. In particular, the domestic unemployment rate proves to be an important determinant of this NTB before and after the onset of the Subprime Crisis. Our paper differs from these insightful studies in an important dimension: There is no need to notify the stricter enforcement of product standards to the WTO.

The theoretical literature mainly views differing product standards as protectionist, since minimum quality standards increase the compliance costs for foreign firms relative to domestic ones. Fischer and Serra (2000) argue that standards chosen by a domestic social planner are always protectionist. In a similar framework, Marette and Beghin (2010) show that domestic standards are not necessarily protectionist, if domestic and foreign producers differ in meeting these costs. However, this only holds for foreign producers being more efficient, an unlikely assumption for developing countries exporting to the United States. Essaji (2010) is interested in the interplay between trade liberalization and the use of

[8] In another study, based on the Global Trade Alert (GTA), Boffa and Olarreaga (2012) conclude that countries have not retaliated during the Great Recession.

product standards. Sturm (2006) offers a political economy model, where uncertainty about the optimal safety level might open the door for hidden domestic transfers.

Our paper is obviously also related to and builds on the empirical literature on product standards.[9] Moenius (2004) provides an important early account on the effects of standards on trade between OECD countries at the industry-level. Several studies exploit the number of notifications of newly imposed product standards by importing countries under WTO's Sanitary and Phytosanitary (SPS) and Technical Barriers to Trade (TBT) agreements or counter-notifications under the SPS for a cross-section or panel of trade flows (see Disdier et al., 2008; Crivelli and Gröschl, 2016; Fontagne et al., 2015). Crivelli and Gröschl (2016) find for a disaggregated gravity model for agricultural and food products that SPS measures decrease the probability of market entry, but positively influence the intensive margin of exporters. For a rich panel data set of French exporting firms from 1996 to 2005, Fontagne et al. (2015) show that restrictive SPS measures (as captured by specific trade concerns (STCs) raised at WTO's SPS Committee) in the importing country negatively affect the extensive and, in contrast to Crivelli and Gröschl (2016), also the intensive margin of trade.

3. Background

3.1. The U.S. Food and Drug Administration (FDA)

The U.S. Food and Drug Administration (FDA) is located within the U.S. Department of Health and Human Services.[10] The FDA is responsible for enforcing the Federal Food, Drug, and Cosmetic (FD&C) Act of 1938 and other laws designed to protect consumer health. The following product categories fall under FDA jurisdiction: food, drugs, cosmetics, medical devices, electronic items that emit radiation, vaccines, blood and biologics, animal feed and veterinary, and tobacco products. To ensure that products from these categories comply with U.S. product standards, the FDA has the authority to inspect domestically produced and imported products and eventually refuse entry into U.S. markets. An inspected domestic

[9] We focus here on the most closely related papers. Other studies differ in their measurement of product standards, the sectors and time covered as well as their approach to the endogeneity issue. The earlier literature focuses on one particular standard in a given product-group (see for instance Otsuki et al., 2001, on African groundnut exports to Europe; Anders and Caswell, 2009, on U.S. seafood imports; Maertens and Swinnen, 2009, on vegetable exports from Senegal; Baylis et al., 2010, on seafood exports to the EU).

[10] For an excellent overview of the FDA, see for instance Buzby et al. (2008), Josling et al. (2004), Hawthorne (2005) and Liu (2010).

product is refused entry if it violates U.S. product standards. However, an imported product can already be refused entry "if it [only] appears to violate" a certain U.S. product standard (Buzby et al., 2008; Liu, 2010). This formulation in Section 801(a) of the Federal Food, Drug, and Cosmetic (FD&C) Act leaves room for discriminatory action of FDA officials with respect to imports.[11] The FDA separates violations into two main categories: adulteration and misbranding. According to the FD&C Act, adulteration means that due to the addition of a substance a product is inferior, impure and not genuine. Most violations for adulteration deal with safety, packaging integrity or sanitation, but differing product standards between trading partners might also be the cause.[12] Besides adulteration, a product might also be denied entry in the United States due to misbranding. Misbranding includes untruthful or misleading statements on product labels or products missing appropriate labeling or packaging (Buzby et al., 2008). This category also comprises products that were rejected by the FDA due to the lack of necessary certificates for conformity assessment.

According to the FD&C Act, every importer of an FDA-regulated product has to file an entry notice with U.S. Customs and Border Protection (CBP), which then notifies the FDA of the entry. The import requests are collected and processed by the computer-based system "Operational and Administrative System for Import Support (OASIS)". The FDA uses OASIS to review the entry documents and to make admissibility decisions. If the FDA does not wish to inspect the entry, the product will proceed into U.S. commerce. If the FDA decides to examine the entry, the importer will not be allowed to further distribute the shipment until the result of the inspection is received. Two types of inspections exist: field exams and sample analysis involving a laboratory test of product samples. The overwhelming majority of inspections are field exams at the ports of entry, whereby FDA inspection officers mainly use organoleptic testing (e.g., appearance and smell) to decide whether a product complies with U.S. product standards.[13] If the product appears to violate these standards, the importer will be given the opportunity to submit a petition to recondition the product into compliance (Buzby et al., 2008; Liu, 2010; FDA, 2011a).[14]

[11] Imported products can be refused without any physical evidence, e.g. just on the basis of bad reputation due to past events of non-compliance at the firm or country-product level (see for instance Jouanjean et al., 2012).
[12] Non-food products can also be refused due to adulteration, i.e., if product-specific regulations are not met.
[13] Barrionuevo, Alexei "Food Imports Often Escape Scrutiny," The New York Times, May 1st, 2007.
[14] Many law firms in the U.S. are specialized on contesting FDA decision of detentions and refusals (e.g., FDAimports.com, LLC: http://www.fdaimports.com/). The services of these law firms are expensive and it is hard(er) for exporters from developing countries to cover such legal costs.

Based on OASIS, the FDA collects information on all ultimately refused shipments in the Import Refusals Report (IRR). The IRR database is available from the beginning of 2002 onwards and includes the exact date of the refusal, name, address and country of origin of the exporting firm, an FDA-specific product code and the product description, port of entry, reason for the refusal and the type of inspection. The database does not include information on the quantity, weight or value of refused shipments, but it is the best source of information on import refusals due to non-compliance with U.S. product standards.

It is important to bear in mind that the FDA´s decision to inspect an entry is not random. The FDA is only able to inspect about 1% of all imported products under its jurisdiction (Buzby et al., 2008; FDA, 2010). To economize its resources for inspections, the FDA employs risk-based criteria to guide its inspections. Using the OASIS database and past import refusals, the FDA identifies exporting countries, product-groups, products or certain firms that have a higher risk of violating U.S. product standards. To react to urgent risks, the FDA additionally issues import alerts that place a product from a certain country or a particular firm on detention without physical examination. Thus, subsequent shipments from this company or country-product-group will be refused automatically, unless the importer can present evidence to overcome this violation.[15] The FDA may also use external information to identify risk products such as the information from the EU-RASFF authorities (Jouanjean et al., 2012). Import surges in a given country-product-group can also trigger more inspections, since any non-compliance represents a higher risk for U.S. citizens. Another reason for an increase in inspections of country-product-groups with higher imports may be protectionism (Trefler, 1993; Essaji, 2008; Baylis et al., 2009).

The United States is an important export market for many countries and about 20% of the overall U.S. imports fall under the jurisdiction of the FDA.[16] 25 cents of every dollar spent on commodities by U.S. consumers are for products regulated by the FDA (FDA, 2011b). A growing share of these products comes from developing countries.[17] In 2010, 15% of food products, 28% of drugs and 52% of medical devices sold in the U.S. markets were

[15] It can be quite costly, in particular for exporters from developing countries, to obtain the necessary documents for conformity assessment from accepted certification bodies (Jaffee and Henson, 2005).

[16] The FDA estimates this share to be over 10%, but only considers food, drugs and cosmetics (FDA, 2011b). Hence, our estimate of around 20% in the year 2011 also includes medical devices, electronic items emitting radiation, animal feed and animal drugs and biologics under the jurisdiction of the FDA.

[17] Emerging markets like China, India and Mexico have increased their exports in FDA regulated products to the U.S. significantly in the last years. Drugs, medical devices and electronic items emitting radiation are the product categories that have experienced the strongest rise in imports from developing countries (FDA, 2011b).

imported. Import lines of FDA regulated products have grown from 6 million in 2001 to 24 million in 2011, corresponding to a 15% annual increase (FDA, 2011b). Note that the resources dedicated to the FDA and the funding provided for FDA officers in the field (who are responsible for product inspections) vary over time (Graph 6). In the aftermath of the September 11 terrorist attacks in 2001 and the Subprime crisis in 2008, the U.S. Congress granted more authority and additional resources to the FDA. The majority of the FDA investigators are assigned to inspect domestic products and facilities.[18] It is difficult to identify the number of FDA officers assigned to the border from official documents.

3.2. Descriptive Statistics on U.S. Imports and Import Refusals

Graph 2 shows total U.S. imports in FDA regulated products and the total number of refused shipments (total refusals) for the years 2002 to 2014. Except for the Great Recession, the overall volume of U.S. imports in FDA regulated products has steadily increased and more than doubled during the sample period.[19] In contrast to imports, import refusals exhibit more variation over time. Graphs 3 and 4 allow for a comparison between OECD and non-OECD countries. There are two main takeaways. First, both country groups share a similar growth pattern in imports, but non-OECD countries account for on average USD 250 billion or about twice the overall import volume of OECD-countries. Second, while both groups had to face an increase in import refusals after the Subprime crisis, the strong co-movement between the U.S. unemployment rate and import refusals from 2008 onwards is clearly driven by poorer countries, whose increase in import refusals is more pronounced and who start from a higher level of total refusals (around 10,000 vs. 4,000 refusals in the year 2009).

To shed more light on the distribution of U.S. import refusals across product-groups, Table A1 in the Appendix shows import refusals at a more disaggregated product level. Food products play a prominent role among those products most often refused during the sample period. Fish products, fruits and vegetables, sugar confectionary, bread and pastry as well as sauces, mixed dressings and condiments are among the top ten most refused product-

[18] Racino, Brad (2011), "Inspectors Struggle to Keep Up with Flood of Imports," News 21 (http://foodsafety.news21.com/2011/imports/border/; download on October 29th, 2014).

[19] Note that the shares for five aggregated product-groups have been quite stable over time (see Online Appendix 1), i.e., the growth in imports is fairly spread over different sectors. Furthermore, the non-food product-groups pharmaceuticals, cosmetics and other manufacturing goods combine for about 75% of total imports in FDA regulated products and are responsible for an increase in total import refusals during the sample period (see Online Appendix 2).

groups. However, the two product categories with most import refusals are other drugs and medical devices. Table A2 in the Appendix includes the ten most frequent reasons for import refusals from 2002 until 2014, showing that import refusals due to adulteration are less frequent than refusals due to misbranding or missing certifications.

To emphasize the importance of FDA regulated products for countries exporting to the U.S., we compute the share of FDA regulated products in total exports to the U.S. for 2012. Graph 5 shows that, for most countries, FDA regulated products comprise more than 20% of total exports to the United States, whereby for some developed and developing countries like Ireland or Denmark and Ghana or Thailand this share rises above 50%.[20]

4. Data and Empirical Strategy

4.1. Data

This paper is based on a newly collected data set. We carefully gather detailed information from two main data sources. Since the FDA uses its own unique product classification system, the main challenge has been to combine the FDA's Import Refusals Report (IRR) database with disaggregated international trade data (c.i.f.) as provided by the U.S. International Trade Commission (ITC). The raw data provided by the FDA reports incidences of import refusals at the firm- and product-level. We aggregate import refusals to the most fine-grained product-group for which a consistent match between the FDA and the Harmonized System (HS) classification is possible. Our guiding principle for this careful matching procedure has been that any FDA product code uniquely falls into the assigned HS product categories. We have succeeded in matching all FDA regulated products to the corresponding HS categories. The exact mapping for our 93 food and non-food product-groups is documented in the table of Online Appendix 5.[21, 22]

[20] Online Appendix 3 and 4 provide more details on the underlying FDA regulated exports for all OECD and non-OECD countries (top 10 exporters of FDA regulated products by region).

[21] In principle, we match FDA product codes to HS 4 digit codes and preserve as much detailed information as possible. For some matched groups, we have to use additional HS 5 or HS 6 digit information. We have succeeded in creating 93 matched product-groups. Note that for medical devices and radiation emitting products, the constructed product-groups may include more HS products than necessary for matching, because it was not possible to isolate FDA regulated products at the HS 6 digit or even at the HS 10 digit classification.

[22] Note that our paper shares a common caveat with the literature that empirically investigates the effects of tariffs and non-tariff barriers on trade flows, namely a possible aggregation bias. As Anderson and van Wincoop (2004) point out, „matching up tariff, nontariff and trade-flow data requires aggregation, guided by concordances that are imperfect and necessarily generate measurement error." Furthermore, Hallak (2010)

Our panel data set starts with the first year for which the IRR data is available, covering on a yearly basis all country-product-groups with at least one notified refusal during the sample period 2002 to 2014.[23] For an important extension of our baseline regression model, we draw on an additional data source. We use notifications from the Rapid Alert System for Food and Feed (RASFF) database of the EU to instrument for U.S. import refusals.[24] The RASFF notification data, which is publicly available at the website of the European Commission, does not include any product identifier. However, the data kindly provided by the German Federal Office of Consumer Protection and Food Safety includes verbal product identifiers, which we carefully match to our 93 product-groups. Furthermore, we obtain disaggregated HS trade flow data for all countries of the world from the World Integrated Trade Solution (WITS) database. We construct a measure first developed by Vollrath (1991) and proposed by Essaji (2008) to compare the import demand structure between the U.S. and other middle and high income countries. This measure of relative import advantage (RMA) allows us to evaluate, whether notifications from EU member countries constitute a valid external instrument.

Finally, we gather a rich dataset on all notifications concerning the WTO's SPS and TBT agreements from WTO's Integrated Trade Intelligence Portal (ITIP). Under these agreements, WTO member countries are required to notify new regulatory measures that may have a "significant effect on trade." Using information on the HS 4 digit sectors affected by the notified measures, we are able to match all notifications made by the United States to the WTO with our 93 product groups. This allows us to control for changes in U.S. product standards regulation during our sample period.

4.2. Empirical Strategy

We proceed in three steps. We start with OLS and standard fixed effects estimates. Then, we follow Arellano and Bond (1991) and estimate a dynamic panel model. The Arellano-Bond

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shows that controlling for determinants of trade at the SITC 2 digit level considerably lowers the aggregation bias in a sectoral gravity estimation. We control in all our main regressions for country-product-group fixed effects and our 93 matched product-groups correspond to the HS 4 digit level (which is considerably more disaggregated than the SITC 2 digit level). Note that about half of our product-groups consist of exactly one HS 4 digit code. This in conjecture with the fact that no single product-group dominates the import refusals coefficient (see Graph A2) makes a strong aggregation bias in our context very unlikely.

[23] Note that we have updated the dataset from our working paper version for the years 2013 and 2014.

[24] We use all the information available in the RASFF notification database (i.e., import refusals and information on detentions and import alerts) to construct our instrument. Jouanjean et al. (2012) acknowledge that these types of information are relevant for inspection authorities at the U.S. FDA.

estimator is a natural choice against the background of large N and small T, a dynamic data generating process and concerns about potential endogeneity. Our bilateral gravity model for disaggregated import flows to the United States covers up to 93 product-groups per country for the years 2002 to 2014, with 170 exporting countries entering our baseline regression. We estimate the following reduced form model:

$$y_{i,k,t} = \sum_{s=1}^{S} (\alpha_s y_{i,k,t-s}) + \beta X_{i,k,t-1} + \delta_t + \mu_{i,k} + u_{i,k,t}, \qquad (1)$$

whereby the dependent variable $Y_{i,k,t}$ measures the real value of imports (c.i.f., in logarithm) from country *i*'s product-group *k* at time *t* to the United States. Our main variable of interest, import refusals, is based on incidences of non-compliance with U.S. product standards for country *i* in product group *k* and year *t* and enters equation (1) with the first lag, $X_{i,k,t-1}$. Following the literature, our preferred measure for import refusals is the following: We use the log(1+refusals) in order to account for the intensity of import refusals, i.e., the total number of refusals in such a country-product-group in time *t*.[25] We control in all our main specifications for time fixed effects (δ_t) and for country-product-group fixed effects ($\mu_{i,k}$).

The Arellano-Bond GMM estimator allows for using internal and external instruments. On the one hand, we instrument the predetermined import refusals with lagged refusals as internal instruments. In our context, this strategy seems very sensible, since FDA inspections are inter alia guided by past incidences of refusals. On the other hand, we use notifications and refusals from the EU-RASFF database as an external instrument for U.S. import refusals.[26] In another important extension, we incorporate WTO notifications (as an additional strictly exogenous variable)[27] to control for legal changes in U.S. product standard regulation. [28]

We will see below that the valid instruments for predetermined and exogenous variables differ. This augmented baseline regression takes the following form:

[25] We also provide results for an alternative measure, i.e., a dummy variable which takes the value of one, if in a given product-group *k* from country *i* at least one incidence of a refusal has been recorded at time *t* (see Table A3 in the Appendix).
[26] We use three types of instruments. First, we use all notifications (information, warnings and import refusals) and, second, only warnings and import refusals from the RASFF-member states Finland, Germany, Norway and Sweden. As a third specification, we include warnings and import refusals from Germany. We generate our variables for non-compliance with EU countries' product standards analogously to the U.S. ones.
[27] Note that strictly exogenous variables are uncorrelated with current and past errors.
[28] One might also argue that WTO notifications, representing legal changes in U.S. product standard regulation, should be assumed to be predetermined. Our results are robust to this alternative assumption.

$$y_{i,k,t} = \sum_{s=1}^{S} (\alpha_s y_{i,k,t-s}) + \beta X_{i,k,t-1} + \gamma Z_{i,k,t-1} + \delta_t + \mu_{i,k} + u_{i,k,t}, \qquad (2)$$

whereby the variable $Z_{i,k,t-1}$ stands, in this case, for the additional external instrument, entering the equation in the same time period as the import refusals, with the first lag. From the differenced form of equation (2), it becomes apparent that the individual fixed effects are eliminated and that the Arellano Bond estimator only uses variations within the country-product-group over time for identification. To clarify our estimation strategy, we continue by using a (simplified) version of this extended regression equation (2).[29]

$$W'\Delta y = W' \sum_{s=1}^{s=2} (\Delta y_{t-s}) \cdot \alpha_s + W'(\Delta X_{t-1}) \cdot \beta + W'(\Delta Z_{t-1}) + W'\Delta u, \qquad (3)$$

whereby a GLS-estimation of equation (3) leads to the one-step Arellano-Bond estimator (not shown). More importantly, this differenced equation is pre-multiplied with matrix $W_{i,k}$:

$$W_{i,k} = \begin{bmatrix} y_{i1}, x'_{i,1}, z'_{i,1}, z'_{i,2}, z'_{i,3} & & & 0 \\ & y_{i,1}, y_{i,2}, x'_{i,1}, x'_{i,2}, z'_{i,1}, z'_{i,2}, z'_{i,3}, z'_{i,4} & & \\ & & \ddots & \\ 0 & & & y_{i,1},.., y_{i,T-3}, x'_{i,1},.., x'_{i,T-3}, z'_{i,1},.., z'_{i,T-1} \end{bmatrix} \qquad (4)$$

The strictly exogenous (here lagged) variable enters contemporaneously in the matrix $W_{i,k}$. For our main variable of interest, the predetermined lagged import refusals, we use its second through fourth lag as internal instruments. To foresee important specification tests of our dynamic panel estimates, the null hypothesis of the validity of the over-identifying restrictions of the Sargan and Hansen tests will indeed never be rejected. Furthermore, the null hypothesis of under-identification and weak identification (overall and for individual endogenous variables) are rejected for (almost) all main regressions.

Note that the time fixed effects capture time-varying characteristics of the importing country, global macroeconomic conditions and factors affecting trade costs for all exporting countries to the United States alike. This time fixed effect also absorbs any changes in FDA inspection capacity in the United States. Furthermore, the country-product-group fixed effects control for the time average of the multilateral resistance terms at the country-

[29] For readability, we omit the time-fixed effects and the indexation for the group dimension. For further details on this GMM estimator, see also Arellano and Bond (1991), Roodman (2009) and Baltagi (2008), whose notation we are largely following in this section.

product-group level and time invariant country-product-group characteristics, like trade costs or production levels. Since the country-fixed effects are a linear combination of these country-product-group specific effects, we are not able to either include country dummies or distance to the United States (a classical gravity variable) separately in the regression.[30]

We continue the discussion of our empirical strategy by outlining our external instrument. First, notifications from the RASFF database constitute a relevant instrument for U.S. import refusals, since U.S. FDA agents are reported to also use external information to identify high-risk products to guide their inspections. It is reassuring for our empirical strategy that Baylis et al. (2009) show that EU import refusals are indeed one important determinant of U.S. import refusals. Second, for the instrument exogeneity to be fulfilled, the same motives for protectionism that may drive U.S. import refusals should not simultaneously influence the notifications from the EU-RASFF database. In contrast to the import refusals from the U.S. FDA, the notifications in the EU-RASFF database are collected from single member countries and do not originate from one central inspection authority. Each country participating in RASFF (EU 28, Norway, Iceland, Switzerland and Liechtenstein) commands its own national food safety system and has the exclusive jurisdiction over food safety inspections and rejections of non-complying import shipments within its borders. Since the countries participating in RASFF strongly differ in their import demand structures, it is very unlikely that the overall EU-RASFF notifications and refusals are jointly driven by protectionist motives.

However, to investigate, whether the import patterns between the U.S. and RASFF member countries are largely uncorrelated, we construct a measure first developed by Vollrath (1991) and proposed by Essaji (2008) to compare the import demand structures between the U.S. and other middle and high income countries. This measure of relative import advantage (RMA) is computed for each product category within each country. The RMA measure takes a value greater (smaller) than 1, if the country imports relatively more (less) of the product than the world average. Essaji (2008) states that the correlation of the RMA measure of all products between two countries is a good measure for the degree of

[30] Note that we will report robustness checks below controlling for time-varying country dummies (multilateral resistance terms) and time-varying aggregated sector dummies.

similarity of import patterns.[31] Graph 10 shows the correlations of the RMA measure between the U.S. and many high and middle income countries using our product groups (average for sample years 2002 to 2014). The correlation between the U.S. and European import patterns (as exemplified by important European countries like Germany or the Nordic countries Finland, Norway and Sweden) is close to zero. This additional piece of evidence is clearly in favor of using notifications and import refusals from the EU RASFF database as external instrument for U.S. import refusals.

However, as an additional robustness check, we exclusively use refusals and warnings notified by Germany as an external instrument. The main argument is that German and U.S. import patterns are uncorrelated (Graph 10). Another important argument for using German notifications as an external instrument is that the German food safety system is particularly transparent and includes many checks and balances to prevent interference in the inspection procedure due to potential protectionist motives. As detailed in the country profile for Germany prepared by the European Commission, a strict rotation system of inspection officers between inspected product sectors exists to prevent possible conflicts of interest (FVO 2013). The two-pairs-of-eyes principle during inspections and the requirement to declare any secondary functions that might lead to possible conflicts of interest are other key measures to ensure impartiality of inspectors.

5. Empirical Results

5.1. Baseline Results

We now turn to our main results. Table 1 presents the estimates for a number of baseline regressions, whereby the refusal intensity log(1+refusals) is our measure for non-compliance with U.S. product standards.[32]

[31] The trade flow data in our paper is based on the HS-classification. To be consistent with our empirical analysis, we compute the RMA measure using our product classification that is in concordance to the HS product classification. Thereby, we define total imports of a country as total imports in FDA regulated products. However, to be able to compare our results to Essaji (2008), we additionally follow his approach and also compute the RMA measure using SITC 4 digit level data. Because our study uses a panel from 2002 until 2014, we compute the RMA measures for each year from 2002 until 2014 as well as for the average trade flows during this time period. Results for RASFF member countries do not change and can be obtained from the authors.

[32] The empirical results for an alternative measure for non-compliance with U.S. product standards (a dummy variable) are reported in Table A3 in the Appendix. We find the same empirical pattern as in Table 1.

[Table 1 about here]

We benchmark our preferred dynamic panel model in Column (4) with pooled OLS and simple fixed effects estimates in Columns (1) and (2), respectively. Compared to Column (4), Column (3) takes the allegedly wrong assumption that refusals are exogenous. Our preferred specification in Column (4) is based on the two-step Arellano-Bond estimator, whereby the dependent variable enters with its first and second lag. This second lag ensures that the Arellano-Bond test for no second-order autocorrelation in first-difference residuals is not rejected, a prerequisite for this estimator to be consistent.[33] We instrument these lagged dependent variables with the first through third lags. Furthermore, for our lagged refusal indicator variable, we use its second through fourth lag as internal instruments. To avoid weak instrument problems, we reduce the number of instruments by collapsing the instrument matrix (Roodman, 2009). All our main results are robust to the exact lag length of our instruments (as shown in the Online Appendix 6 and 7).

There are two main findings from Table 1. First, the disaggregated U.S. import flows exhibit substantial persistence over time. Our estimates show that the first and the second lag of the dependent variable enter our regression significantly. This corroborates our decision to use a dynamic panel estimator. The null hypothesis of instrument validity is not rejected for any of the two tests of overidentifying restrictions, the Sargan or Hansen test, in any main regression. Furthermore, additional specification tests show that we can also reject the null hypotheses of under-identification and weak instruments (see section 5.5. and Table A5 for further details).[34] All these tests indicate that our preferred specification is well-specified.

Second, import refusals are indeed endogenous to import flows. When we control for lagged import flows and country-product-group fixed effects using the Arellano-Bond estimator in Column (3), the positive and significant impact of refusals on import flows of the

[33] The Arellano-Bond test for no first-order autocorrelation is rejected, as we expect from econometric theory.

[34] In particular, we employ the tests for under-identification and for weak instruments as proposed by Sanderson and Windmeijer (2016), Kleibergen and Paap (2006) and Stock and Yogo (2005) as well as Yogo (2004), respectively. In the latter case, we follow Bazzi and Clemens (2013) and test for the null hypothesis that the bias is greater than 30 percent of the OLS bias.

simple OLS in Column (1) disappears. Once we additionally instrument our import refusals indicator with lagged refusals, the point coefficient becomes negative and significantly different from zero at the 10% level (see Column (4) of Table 1). These results demonstrate that the endogeneity of trade barriers leads to a strong upward bias in the estimated coefficient and an underestimation of trade costs, if the empirical strategy does not address these endogeneity issues. Trefler (1993) finds that accounting for endogeneity of non-tariff barriers to trade (NTBs), the estimated negative impact of NTBs on U.S. imports is ten times larger than in estimations not addressing the endogeneity issue. Similar to our paper, Essaji (2008) shows that the effects of U.S. technical regulations on U.S. imports are significantly negative when accounting for endogeneity of technical regulations to import flows. In the case of not addressing the endogeneity issues, he reports positive effects of technical regulations on imports.

5.2. Developing vs. Developed Countries

In this section, we go beyond the overall impact of import refusals for all countries and present a more nuanced picture by focusing on non-OECD countries and by distinguishing between different product-groups and types of refusals for OECD and non-OECD countries.

We start by restricting the baseline regressions of Table 1 to non-OECD countries only in Table 2. While OECD-countries include all industrialized countries with a very high standard of living, non-OECD countries encompass developing and emerging markets with on average a lower GDP per capita.[35] The empirical results from Table 2 provide first evidence that poorer countries are bearing the trade costs from non-compliance with U.S. product standards, with the point coefficient for refusals (-0.204) being significant at the 5% confidence level. Unreported results show that we find no significant effects for OECD countries.

[Table 2 about here]

[35] Note that we employ the "classical" definition of OECD countries from the beginning of the 1990s (before emerging markets like Mexico and Chile joined), since it provides a sharper distinction between rich and poor countries. None of our main results hinges on the exact definition.

In the first column of Table 3, we report our preferred specification for all countries of Table 1, Column (4), for comparability. We refer to this estimation as our baseline estimation. Column (2) allows the slope coefficient for import refusals to vary by product-group. We distinguish between food-products and non-food products and find a negative and significant coefficient for food-products.

[Table 3 about here]

In Columns (3) and (4) of Table 3, we investigate, whether there is a differential effect of import refusals between developed and developing countries for food and non-food products. Note that we apply the same specification as in Column (2) once to OECD countries and once to non-OECD countries in Columns (3) and (4), respectively. This sample split reveals that the negative impact of refusals on food imports is driven by poorer countries. For non-OECD countries, the point coefficients for food and non-food products are both negative and in the same ballpark, but (due to the larger standard error for non-food products) only the former coefficient is also significantly different from zero.

Furthermore, we offer results for different types of non-compliance in Table 4. We group refusals into refusals due to adulteration and refusals due to misbranding and allow the slope coefficient for refusals to vary by refusal type, but do not find a robust empirical pattern along this dimension.

[Table 4 about here]

5.3. Is there Evidence for Protectionism?

To recap our empirical results thus far: We find that import refusals negatively affect disaggregated trade flows to the United States, in particular for non-OECD countries. But our empirical analyses have not offered any indication for hidden protectionism yet.

We will argue in this section that the type of inspection can shed some light on hidden protectionism. Remember that an imported product can be refused entry, if it simply "appears to violate" U.S. product standards. Hence, there is considerable leeway for the FDA to enforce these standards, opening the door for less honorable motives than pure health or product quality concerns. It is reasonable to assume that this leeway is most pronounced, if a refusal is not based on any laboratory tests but solely on the judgment of an FDA officer. We will proceed in two steps in this section. First, we will distinguish between refusals with and without any product sample analysis. Second, we are interested in examining to what extent refusals and the negative trade effect of the inspection type varies over time.

Table 5 presents the results for different types of inspection that lead to import refusals. The results in Column (2) indicate that those refusals that are not based on any product sample analysis are driving the negative effect of (the total number of) refusals on imports for the overall sample. Even worse, when comparing the results in Column (3) and (4), it becomes evident that solely non-OECD countries suffer from this discretionary room for refusal decisions. To be clear, many of these refusal decisions might be well-grounded, for instance if a product is obviously rotten. But the room for discriminatory action is considerable in this category and it is worrisome that only developing countries are suffering from these potentially arbitrary refusals.[36] It is exactly this type of non-tariff barrier to trade that might fly under the radar, since it is hard to identify and measure.

[Table 5 about here]

In an intermediate step, Table 6 sheds some light on the evolution of refusals over time. The first column reports a simple OLS regression with two explanatory variables, namely the number of refusals and warnings in a given product-group-year (as reported by the RASFF member countries Germany, Finland, Norway and Sweden) and a dummy variable that takes the value of one, if a given product-group is affected by at least one notification (within the TBT or SPS agreement) of the United States to the WTO in a given year. Interestingly, the first variable, RASFF-refusals, only significantly explains U.S. import refusals in the United States, when we do not account for product-group fixed effects (starting from

[36] Various cases of discriminatory action of U.S. authorities against imports from developing countries are documented in the literature. For instance, U.S. authorities have banned Mexican Hass Avocados from entering the U.S. market for 79 years due to pest concerns, though officials from the U.S. Department of Agriculture have repeatedly certified Mexican growing areas as pest free during that time period (Lamb, 2006).

column 2). While the simple fixed effects estimates in Columns (2) to (6) should not be interpreted causally, they clearly indicate a positive correlation between the number of refusals and the Subprime Crisis and its aftermath in the United States. This also holds true, when we control for de jure changes in U.S. product standards (WTO notifications) during the sample period. Hence, the spike in refusals during the time period 2008 to 2012 does not seem to be driven by stricter product standards. Finally, we learn from the comparison between Columns (5) and (6) that only non-OECD countries are facing significantly more import refusals during crisis times. This significant rise in refusals by 4 to 6 percent is robust to the inclusion of two potentially endogenous variables, namely the lagged value of import refusals and (in unreported results) imports (in logarithm). This provides further evidence consistent with hidden protectionism during crisis times.[37] Our argument is further supported by Levchenko et al. (2011), who show that the product quality of U.S. imports did not decline during the Subprime Crisis and its aftermath. Thus, the significant rise of import refusals against non-OECD countries cannot be explained by lower product quality of non-OECD exports.

[Table 6 about here]

Finally, we show in Table 7 how the results for both types of inspection for OECD and non-OECD countries vary over time. For comparability, Columns (1) and (3) report the baseline results for both country groups for refusals based on product sample analyses and refusals without such analyses. More importantly, we present in Columns (2) and (4) the results for the period from 2008 to 2012 that encompasses the Subprime Crisis and its aftermath in the United States.[38] Note that the average unemployment rate in the United States rose from 5.27 (2002-2007) to 8.34% (2008-2012).

[37] The Subprime Crisis and its aftermath (2008-2012) also coincide with a rise in the unemployment rate and an increase in the number of inspections and of FDA field officers as well as the FDA budget more generally (Graph 6). Since all these variables are highly, positively correlated, it is not sensible to enter them simultaneously in such a regression. Note that these variables only vary over time but not across country-product groups and are captured in all our main regressions by the time fixed effects.

[38] Further results on the type of inspection for 5-year rolling windows for OECD and non-OECD countries are presented in Online Appendix Tables 8 and 9.

[Table 7 about here]

The results of Table 7 are striking. Non-OECD countries generally suffer from import refusals without any product sample analysis and they suffer all the more when unemployment rates in the U.S. are at historical highs. In stark contrast, import refusals do not have any statistically significant impact on export flows for OECD countries to the United States during any time period. In addition, Graph 6 offers another interesting insight. The share of FDA-inspections based on a product sample analysis (out of all FDA-inspections with and without product sample analysis) has decreased over the last few years, even though the total FDA budget for field activities (i.e., product inspections) and the number of FDA-officers in the field have increased. In our view, the regression results are consistent with the hypothesis that the enforcement of U.S. product standards is hidden protectionism in disguise at the cost of poor countries.

Is it possible to identify the biggest losers from U.S. hidden protectionism during this crisis times? Yes, but only indirectly. Graph 7 plots the point coefficients and 90 percent confidence intervals for import refusals (without a sample analysis) for 145 regressions. The very last data point in the graph represents the estimate for non-OECD countries and the years 2008-2012 from Table 7, Column (4). The remaining 144 observations correspond to regressions, in which we consecutively eliminate one of the 144 countries from the sample.[39] Hence, we can infer from the largest decreases in the size of the coefficient those countries that suffered most from import refusals without any sample analysis. The four countries that triggered the three biggest upward spikes are China, Chile, Sri Lanka and Russia. In a similar vein, according to this measure, the Ukraine comes off best.[40] Pure coincidence? Maybe, but diplomatic tensions between the United States and China as well as Russia have markedly increased over the last few years. The U.S.-Russia-relations deteriorated to their lowest-point since the Cold War due to the Russia-Georgia conflict in August 2008 (Nichol, 2014).

[39] In a similar vein, we check the robustness of the results in Table 7 by eliminating one out of the 93 product-groups in the sample. Graph 8 indicates that the results are not driven by any particular product-group.
[40] Note that (judging from the standard errors) this interesting pattern does not lead to statistically different point coefficients for these subsamples.

Finally, hidden protectionism due to stricter enforcement of product standards might fly under the radar to a certain extent, but not completely. In particular, if U.S. enforcement measures are strong and the affected exporters influential in their home country, foreign governments will contact the WTO. We follow Fontagne et al. (2015) and look at the specific trade concerns (STCs) that have been raised against the U.S. within WTO´s SPS and TBT committees. WTO member countries have the opportunity to raise specific trade concerns against regulatory measures (in draft or in force) of importing countries that they perceive as potentially trade restrictive. Fontagne et al. (2015) argue that the overall increase in the cumulative number of HS 4 digit sectors that are affected by an unresolved STC from 2008 to 2009 is consistent with the general perception of increased protectionism after the crisis. We can confirm this pattern for the United States. Graph 9 singles out the cumulative number of unresolved STCs (by complaining countries and affected HS 4 digit sectors) against the United States and shows a rise in trade concerns in the years 2008 to 2012.

5.4. The Costs of Non-Compliance with U.S. Product Standards

The size of the negative effects of U.S. import refusals on U.S. imports is substantial. An increase of the refusal intensity by 10% reduces U.S. imports in a certain country-product-group by 1.3% (see Table 1, Column 4). Trefler (1993) and Essaji (2008) also report large economic effects of U.S. non-tariff barriers to trade. For instance, Trefler (1993) finds that NTBs reduced manufacturing imports to the United States by 24% in 1983.

In our paper, the negative trade effects of import refusals are confined to emerging and developing countries. An increase in the number of refusals by one standard deviation decreases imports to the United States from non-OECD countries by about real USD 175 million (in 2010 terms) per country and product-group (based on Column (4) of Table 2, which presents the baseline specification for non-OECD countries). Since for non-OECD countries on average about 16 product-groups underlie the identification of this effect, the short-run costs of non-compliance amount to around USD 2.8 billion per exporting country. Turning to the long-run effects (based on the long-run coefficient for import refusals $\beta_{LT}=\beta/[1-(\alpha_1+\alpha_2)]$), we find that developing countries lose close to USD 5 billion in export flows to the United States.

Our empirical results confirm and qualify that product standards and their enforcement represent a challenge in particular to developing countries (Essaji, 2008;

Staiger, 2015). The fixed costs to enter a foreign market are higher for producers in developing countries and their production costs are more sensitive to a tightening of product standards and their enforcement.[41,42] Poorer countries often lack the public infrastructure, investment sources and human capital to meet the product standards and conformity assessment requirements of a developed importing country. Hence, product standards and stricter enforcement procedures negatively affect import flows especially for developing countries (Henson and Loader, 2001; Disdier et al., 2008; Essaji, 2008).[43]

5.5. Further Results

We conclude the empirical analyses with a series of robustness checks. We are interested in the sensitivity of our main results to controlling for changes in U.S. product standards regulation and inclusion of further fixed effects. We also provide tests for under-identification and weak instruments. Finally, we discuss alternative hypotheses.

We start by reporting our baseline results for all countries and non-OECD countries for the years 2002-2014 as well as for non-OECD countries during the crisis period for refusals with and without product sample analyses for comparability in Column (1) of Tables A4, A6 and A8 in the Appendix. So far, our empirical results may also be explained by changes in the strictness of U.S. product standards. In Column (2), we incorporate a further variable that controls for de jure changes in U.S. product standards regulation. Note that all WTO member countries are required to notify (under the SPS and TBT agreement) new regulatory measures that may have a "significant effect on trade". We incorporate a dummy variable that takes the value of one for any product-group, if at least one regular or emergency measure (as notified by the United States) went into force in year t, and zero otherwise. To summarize these additional empirical results, legal changes in U.S. product standard regulation seem to impede imports to the United States to a certain extent, but the main effect of import refusals remains robust. This indicates that stricter enforcement of

[41] For instance, Maskus et al. (1999), Jaffe and Henson (2005) and Jaud and Kukenova (2011).

[42] The literature stresses the costs incurred to meet the precise technical regulation (product re-design) and costs for verifying that regulations are met (conformity assessment). Maskus et al. (1999) claim that conformity assessment costs, which relate to the enforcement of product standards in importing countries, pose by far the larger technical barrier to trade for exporting firms in developing countries. They also state that the recognition of conformity assessment certificates leaves room for protectionism.

[43] For the shrimp industry in Bangladesh and Nicaragua, Jaffe and Henson (2005), Cato et al. (2000) and Cato and Lima dos Santos (2003) provide numerical examples for the sizable adjustment costs developing country exporters face due to stricter conformity assessment procedures established by the United States.

given product standards, in particular during the period 2008 to 2012, constitutes a separate channel through which countries might protect their domestic industries.

Turning to Columns (3) and (4) of Tables A4, A6 and A8, we check whether our main results are robust to the inclusion of further fixed effects. Note that these two specifications are based on those countries that export at least 20 (out of 93) product-groups to the United States. This restriction is necessary, because more cross-sectional variation within countries is required for these more demanding specifications. Column (3) adds sector-year fixed effects to the baseline specification in Column (1).[44] This specification is further extended by including country-time fixed effects in column (4), whereby we consider two-year periods in each case. Thus, in the most demanding specification in column (4), time-, country-product-, sector-year and country-time-fixed effects are included at the same time (on top of the WTO notifications). Hence, in the order of fixed effects, this specification controls for macroeconomic shocks or changes in enforcement capacity of the FDA that affects all product-groups in the United States, any time-invariant country-specific product characteristics, any time-varying global factors at a more aggregated product-level and any time-varying characteristics at the country of origin. It is reassuring that all our main results are insensitive to these alternative specifications, with the point estimates being in the same ballpark and significant at least at the 10% level.

We now turn to another central topic. We follow Bazzi and Clemens (2013) and open the “black box” of GMM in order to assess the strength of our instruments. We present the results of this empirical exercise once more for all our main regressions in Tables A5, A7 and A9, whereby the respective baseline regression is shown in the first column. Note that the upper panel always refers to the GMM estimates and, within the same column, the lower panel displays the IV-equivalent estimated via two-stage least squares. Hence, the IV-results for our baseline regressions are presented in Column (5) of Table A5. We confirm an important finding of the Arellano-Bond GMM estimates in the IV-framework. Once more, the Hansen-test of over-identifying restriction indicates no correlation between the instruments and the error term. Furthermore, the additional tests for under-identification and weak instruments confirm the strength of our internal instruments (Kleibergen and Paap, 2006). A new weak instrument test suggested by Sanderson and Windmeijer (2016)

[44] We define 5 aggregated sectors for our 93 product-groups: fish products, fruits and vegetables, other food products and animal feed, pharmaceuticals and cosmetics and other manufacturing products.

even shows that we can reject the null hypothesis of weak instruments separately for our endogenous variables (imports and refusals). These additional specification tests clearly show that there is no need for an additional external instrument in our setting.

Nonetheless, in Columns (2), (3) and (4), we extend our internal instruments by three different external instruments based on EU's RASFF, namely all notifications from the RASFF-member states Finland, Germany, Norway and Sweden, warnings and refusals reported by those four countries as well as warnings and refusals reported by Germany. For the GMM estimation, our main variable of interest, import refusals, stays significant in all cases.. The IV-estimates are also in the same ballpark, but insignificant in about one third of the cases due to a higher standard error. Turning to the specification tests, we can reject the under-identification test for all additional specifications with the external instrument in Table A5, A7 and A9. The same holds true for all tests of weak instruments, except for Table A9 where we can only reject the hypothesis of weak instruments for the Sanderson-Windmeijer test. To conclude, the external instruments tend to weaken the overall instrument strength, but the overall impression of the under-identification and weak instrument tests remains positive.

5.6. Potential alternative explanations

We would like to wrap up the discussion of the empirical results with some alternative hypotheses. 1) Is it all China? Against the background of buoyant exports to the U.S., numerous press reports about contaminated Chinese imports give the impression that the FDA is basically fighting bad shipments from China. We aim for a more comprehensive robustness check. Similar to Graph 7, we report in Graph A1 the point coefficients and 90 percent confidence intervals for 145 regressions, with the baseline for non-OECD countries in Table 2, Column (4) being the last data point. In each of the remaining 144 regressions, we consecutively drop one of the 144 countries from the sample. There is only one country pivotal enough to push the upper confidence line marginally above zero. Surprise, surprise, it's India.[45] 2) Could a drop in the quality of imports have led to a rise in import refusals and a drop in import values during the Subprime Crisis? This alternative hypothesis would be hard to decisively refute within our empirical setting, but Levchenko et al. (2011) show for

[45] The very same empirical exercise for product-groups shows that our results are not driven by a specific product-group in Graph A2.

very detailed HS-products that the product quality of U.S. imports did not decline during the Subprime Crisis and its aftermath. 3) What might be the underlying channel for the increase in FDA-inspections, import refusals and the drop in imports during the crisis? One might argue that the increase in inspections and, allegedly, enforcement of product standards could also be consistent with an efficiency wage hypothesis. FDA officers increase their efforts and avoid shirking during the downturn of the business cycle, when the outside options in the labor market are limited. This hypothesis would go counter the hidden protectionism story, but why should FDA officers target only non-OECD countries in their increased efforts? While we cannot identify the exact channel underlying our main effect, it seems plausible that lobbying efforts by industry representatives are more successful and the FDA more prone to political pressures during crisis times. Bagwell and Staiger (2003) offer a theoretical model for the countercyclical protectionism.

Last but not least, what could be the reason that non-OECD countries are more strongly targeted by hidden protectionism? Generally speaking, harmonization of product standards and technical cooperation is much more elaborated between OECD-countries than between OECD and non-OECD countries. This argument might be relevant for the overall effects, but less so for the asymmetric rise in import refusals during the Subprime Crisis and its aftermath. One might speculate that the United States is less willing to target OECD-countries with hidden protectionism, because these countries could more easily and more forcefully retaliate. It seems also worthwhile to carefully compare once more our results with Bown and Crowley (2013), who clearly show that another NTB, temporary trade barriers, has been less strongly used during the Subprime Crisis than we would have expected from historical U.S. data. But, to the extent that TBTs have been used, Bown and Crowley (2013) and our paper share the same targeted group of countries: (Non-OECD) countries with relatively stronger growth during the global economic downturn. Our results suggest that during the last crisis, the United States might have substituted temporary trade barriers (that need to be notified to the WTO) with stricter enforcement of product standards (that might fly under the radar for a while).

6. Conclusions

This paper assesses the impact of U.S. import refusals on U.S. sectoral import flows for a rich data set of 93 product-groups for 170 trading partners from 2002 to 2014. Our estimates show that non-compliance with U.S. product standards can exhibit substantial trade costs. This trade reducing effect of enforced product standards is driven by non-OECD countries, whereby our empirical results indicate that a one standard deviation increase in refusals reduces short- and long-run exports from developing countries by USD 2.8 to 5 billion. Hence, non-compliance with product standards can be very costly for poorer countries and might hinder their economic development.

Furthermore, we find striking evidence that the intensity of FDA inspections and import refusals as well as the negative effect of import refusals on U.S. imports have increased in the aftermath of the Subprime Crisis. During this time period, the unemployment rates in the United States have also markedly increased, suggesting that there is a business cycle element to the non-compliance of imports with U.S. product standards. Furthermore, by controlling for changes in U.S. products standard regulation, we show that the increase of import refusals is due to a stricter enforcement of existing product standards.

Most importantly, we find that the sharp increase in import refusals is driven by those refusals that are not based on any product sample analysis. It is exactly this sort of inspection that offers most leeway for FDA officers. In many instances these refusals are for sure warranted, but it is puzzling that these types of refusals are counter-cyclical, suggesting that the FDA, like any other U.S. agency, might not be immune to political pressures. We conclude that our empirical results are consistent with the existence of counter-cyclical, hidden protectionism due to non-tariff barriers to trade in the United States. Hence, this paper corroborates worries raised by Baldwin and Evenett (2009) about a rise of murky protectionism.

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Tables and Graphs

Table 1: U.S. Imports and Refusals – Different Estimators

Variables	***OLS***	***Fixed effects***	***Arellano-Bond***	***Arellano-Bond***
	(1)	***(2)***	***(3)***	***(4)***
Log Imports (t-1)	0.729**	0.449**	0.416**	0.419**
	(0.01)	(0.01)	(0.03)	(0.03)
Log Imports (t-2)	0.240**	0.061*	0.058**	0.059**
	(0.01)	(0.01)	(0.02)	(0.01)
Refusals (t-1)	0.032**	0.014°	-0.002	-0.134°
	(0.00)	(0.01)	(0.01)	(0.08)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	No	Yes	Yes	Yes
Refusals endogenous	No	No	No	Yes
AR(1) (*p*-value)			0.000	0.000
AR(2) (*p*-value)			0.711	0.669
Sargan-test (*p*-value)			0.600	0.875
Hansen-test (*p*-value)			0.843	0.950
Number of instruments			15	17
Number of groups			3468	3468
Number of countries	170	170	167	167
Number of observations	33787	33787	29882	29882

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). Robust standard errors are reported in parentheses. The estimates in Column (1) and (2) are based on pooled OLS and (country-product) fixed effects, with standard errors being clustered at the country-level. Columns (3) and (4) employ a two-step Arellano-Bond estimator with robust standard errors. The lagged dependent variable is instrumented with its first through third lag. In Column (3), we define the variable refusals as exogenous. In Column (4), we allow the variable refusals to be endogenous and instrument it with its second through fourth lag. The instrument matrix is collapsed. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table 2: U.S. Imports and Refusals – Different Estimators, Non-OECD Countries

Variables	***OLS***	***Fixed effects***	***Arellano-Bond***	***Arellano-Bond***
	(1)	***(2)***	***(3)***	***(4)***
Log Imports (t-1)	0.720**	0.435**	0.385**	0.386**
	(0.01)	(0.02)	(0.04)	(0.04)
Log Imports (t-2)	0.245**	0.066**	0.057**	0.059**
	(0.01)	(0.01)	(0.02)	(0.02)
Refusals (t-1)	0.037**	0.011	-0.004	-0.204*
	(0.00)	(0.01)	(0.01)	(0.10)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	No	Yes	Yes	Yes
Refusals endogenous	No	No	No	Yes
AR(1) (*p*-value)			0.000	0.000
AR(2) (*p*-value)			0.452	0.393
Sargan-test (*p*-value)			0.891	0.936
Hansen-test (*p*-value)			0.955	0.930
Number of instruments			15	17
Number of groups			2496	2496
Number of countries	147	147	144	144
Number of observations	23708	23708	20839	20839

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). Robust standard errors are reported in parentheses. The estimates in Column (1) and (2) are based on pooled OLS and (country-product) fixed effects, with standard errors being clustered at the country-level. Columns (3) and (4) employ a two-step Arellano-Bond estimator with robust standard errors. The lagged dependent variable is instrumented with its first through third lag. In Column (3), we define the variable refusals as exogenous. In Column (4), we allow the variable refusals to be endogenous and instrument it with its second through fourth lag. The instrument matrix is collapsed. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table 3: U.S. Imports and Refusals – Food vs. Non-Food Products

Variables	*Baseline*	*All countries*	*OECD-countries*	*Non-OECD-countries*
	(1)	*(2)*	*(3)*	*(4)*
Log Imports (t-1)	0.419**	0.419**	0.467**	0.387**
	(0.03)	(0.03)	(0.08)	(0.04)
Log Imports (t-2)	0.059**	0.058**	0.045°	0.059**
	(0.01)	(0.01)	(0.03)	(0.02)
Refusals (t-1)	-0.134°			
	(0.08)			
Refusals (non-Food) (t-1)		-0.011	0.282	-0.204
		(0.19)	(0.26)	(0.22)
Refusals (Food) (t-1)		-0.155*	-0.106	-0.209*
		(0.08)	(0.11)	(0.10)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.669	0.575	0.468	0.365
Sargan-test (*p*-value)	0.875	0.942	0.645	0.975
Hansen-test (*p*-value)	0.950	0.948	0.925	0.957
Number of instruments	17	20	20	20
Number of groups	3468	3468	972	2496
Number of countries	167	167	23	144
Number of observations	29882	29882	9043	20839

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. In Column (1), we report the baseline estimates for all refusals from Table 2, Column (4), for comparability. Column (2) allows for different slope coefficients for refusals for food and non-food product-groups. Columns (3) and (4) report the same specification for OECD-countries and non-OECD-countries, separately. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table 4: U.S. Imports and Refusals – Type of Refusal

Variables	***Baseline***	***All countries***	***OECD-countries***	***Non-OECD-countries***
	(1)	***(2)***	***(3)***	***(4)***
Log Imports (t-1)	0.419**	0.422**	0.497**	0.393**
	(0.03)	(0.03)	(0.08)	(0.04)
Log Imports (t-2)	0.059**	0.057**	0.054°	0.057**
	(0.01)	(0.01)	(0.03)	(0.02)
Refusals (t-1)	-0.134°			
	(0.08)			
Refusals (Misbranding) (t-1)		-0.032	0.240	-0.181
		(0.16)	(0.26)	(0.18)
Refusals (Adulteration) (t-1)		-0.127°	-0.147	-0.136
		(0.08)	(0.14)	(0.09)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.669	0.572	0.579	0.281
Sargan-test (*p*-value)	0.875	0.536	0.460	0.458
Hansen-test (*p*-value)	0.950	0.352	0.969	0.271
Number of instruments	17	20	20	20
Number of groups	3468	3468	972	2496
Number of countries	167	167	23	144
Number of observations	29882	29882	9043	20839

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. In Column (1), we report the baseline estimates for all refusals from Table 2, Column (4), for comparability. Column (2) allows for different slope coefficients for refusal types adulteration and misbranding. Columns (3) and (4) report the same specification for OECD-countries and non-OECD-countries, separately. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table 5: U.S. Imports and Refusals – Type of Inspection

Variables	*Baseline*	*All countries*	*OECD-countries*	*Non-OECD-countries*
	(1)	*(2)*	*(3)*	*(4)*
Log Imports (t-1)	0.419**	0.422**	0.506**	0.391**
	(0.03)	(0.03)	(0.09)	(0.04)
Log Imports (t-2)	0.059**	0.060**	0.053°	0.059**
	(0.01)	(0.01)	(0.03)	(0.02)
Refusals (t-1)	-0.134°			
	(0.08)			
Refusals (Sample) (t-1)		0.121	-0.096	0.154
		(0.14)	(0.26)	(0.16)
Refusals (no Sample) (t-1)		-0.154	0.113	-0.252*
		(0.10)	(0.17)	(0.13)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.669	0.601	0.596	0.289
Sargan-test (*p*-value)	0.875	0.739	0.493	0.836
Hansen-test (*p*-value)	0.950	0.622	0.918	0.697
Number of instruments	17	20	20	20
Number of groups	3468	3468	972	2496
Number of countries	167	167	23	144
Number of observations	29882	29882	9043	20839

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. In Column (1), we report the baseline estimates for all refusals from Table 2, Column (4), for comparability. Column (2) allows for different slope coefficients for the type of inspection leading to a given refusal. We distinguish between refusals without any product sample analysis and refusals after an FDA or private product sample analysis has been provided. Columns (3) and (4) report the same specification for OECD-countries and non-OECD-countries, separately. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table 6: U.S. Refusals and the Subprime Crisis – A First View

Variables	*All countries*	*All countries*	*All countries*	*All countries*	*OECD-countries*	*Non-OECD-countries*
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(4)*
Refusals (RASFF)	0.088°	0.004	0.004	0.005	-0.032	0.023
	(0.05)	(0.02)	(0.02)	(0.01)	(0.03)	(0.02)
WTO Notifications	0.049**	0.004	-0.010	-0.012	-0.012	-0.012
	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)
WTO Notifications (t-1)	0.075**	0.019*	0.011	0.010	0.028°	0.002
	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)
Crisis			0.038**	0.042**	-0.001	0.061**
			(0.01)	(0.01)	(0.02)	(0.01)
Refusals (t-1)				0.092**	0.073**	0.100**
				(0.01)	(0.02)	(0.01)
Time FE	No	No	No	No	No	No
Country-Product FE	No	Yes	Yes	Yes	Yes	Yes
Number of groups	3468	3468	3468	3468	972	2496
Number of countries	167	167	167	167	23	144
Number of observations	29882	29882	29882	29882	9043	20839

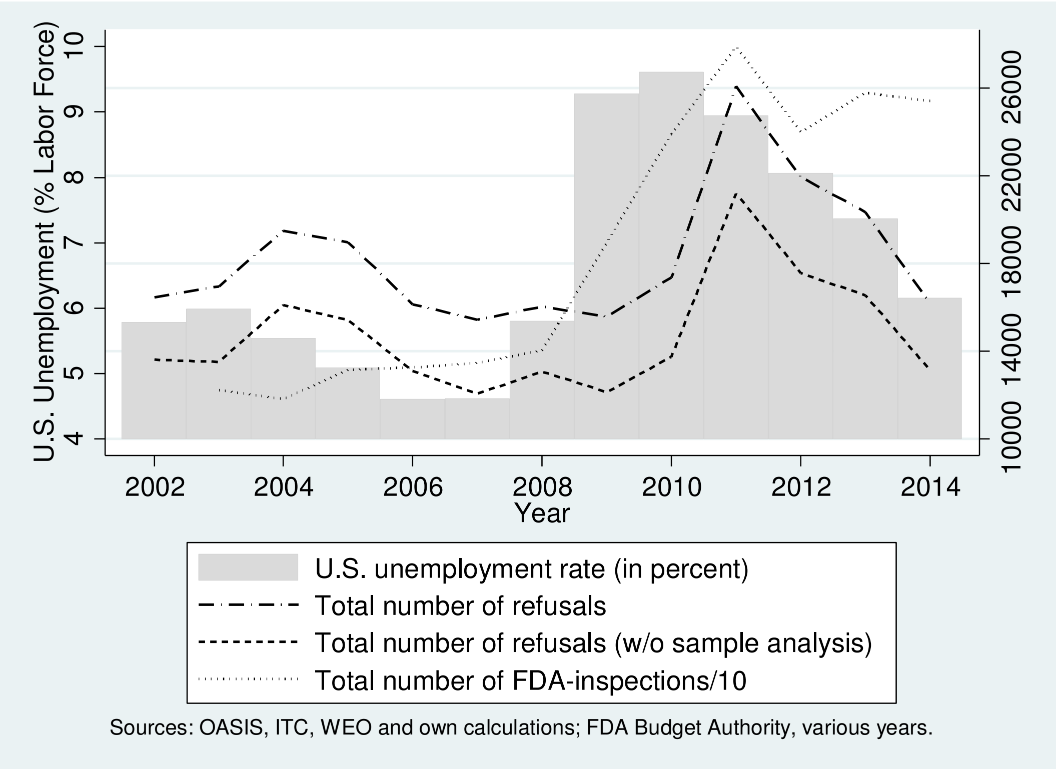
Notes: Dependent variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). The variable Crisis is a dummy variable taking the value of one for the years 2008 to 2012 and zero otherwise. The variable WTO notifications is a dummy variable that takes the value of one if a given product-group is affected by at least one WTO notification (TBT or SPS agreement) of the United States in year t and zero otherwise. Standard errors clustered at the country-product-level are reported in parentheses. The variable Refusals (RASFF) is capturing the number of warnings and refusals reported by the RASFF-members Germany, Finland, Norway and Sweden. All estimates (except column 1) are based on country-product fixed effects, respectively. All estimates are based on the same sample as the main sample (see Table 1). Note that the positive and significant point coefficient on the variable Crisis is robust to the inclusion of the potentially endogenous variable log Imports. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table 7: U.S. Imports and Refusals – Evidence for Hidden Protectionism?

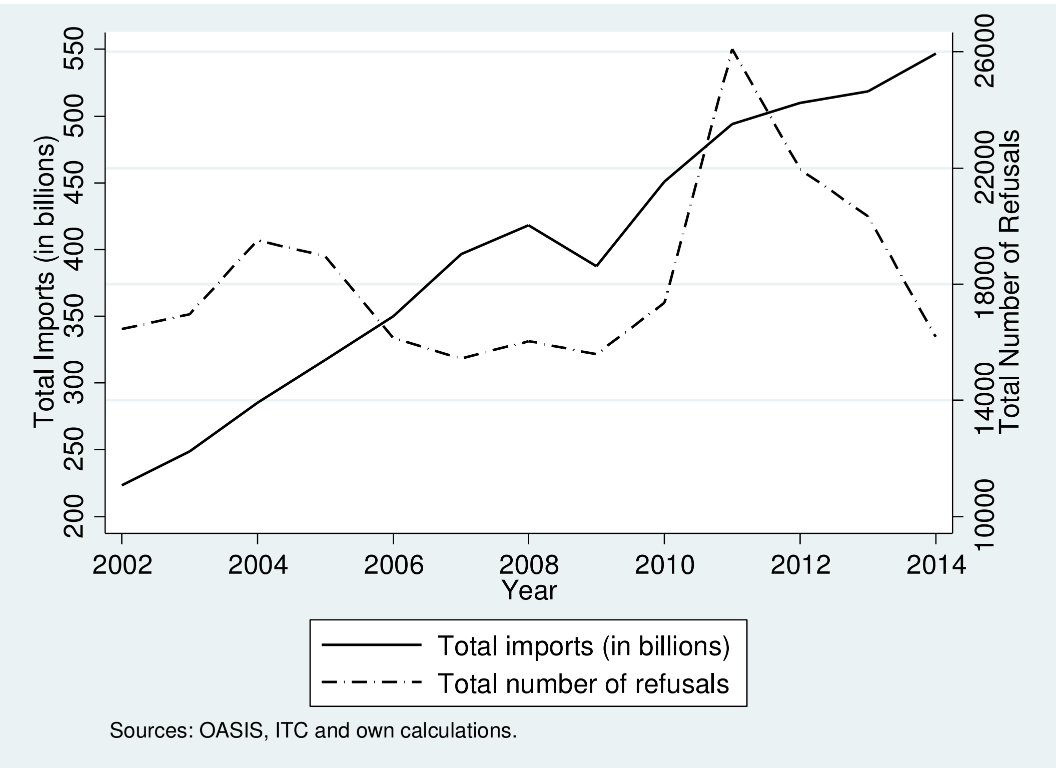
Variables	***OECD-countries***		***Non-OECD-countries***	
	All years	***Crisis years (2008-2012)***	***All years***	***Crisis years (2008-2012)***
	(1)	***(2)***	***(3)***	***(4)***
Log Imports (t-1)	0.506**	0.447**	0.391**	0.385**
	(0.09)	(0.12)	(0.04)	(0.07)
Log Imports (t-2)	0.053°	-0.000	0.059**	0.067*
	(0.03)	(0.04)	(0.02)	(0.03)
Refusals (Sample) (t-1)	-0.096	-0.393	0.154	0.326
	(0.26)	(0.44)	(0.16)	(0.26)
Refusals (no Sample) (t-1)	0.113	-0.075	-0.252*	-0.532*
	(0.17)	(0.16)	(0.13)	(0.26)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.596	0.520	0.289	0.311
Sargan-test (*p*-value)	0.493	0.364	0.836	0.911
Hansen-test (*p*-value)	0.918	0.960	0.697	0.931
Number of instruments	20	15	20	15
Number of groups	972	942	2496	2334
Number of countries	23	23	144	144
Number of observations	9043	4533	20839	10491

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. In Column (1) and Column (3), we report the estimates for the whole time period from Table 5, Columns (3) and (4) for comparability. Note that results in columns (2) and (4) of Table 7 refer to the "crisis years" from 2008 to 2012, a time period dominated by the Subprime Crisis and rising unemployment rates in the United States. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

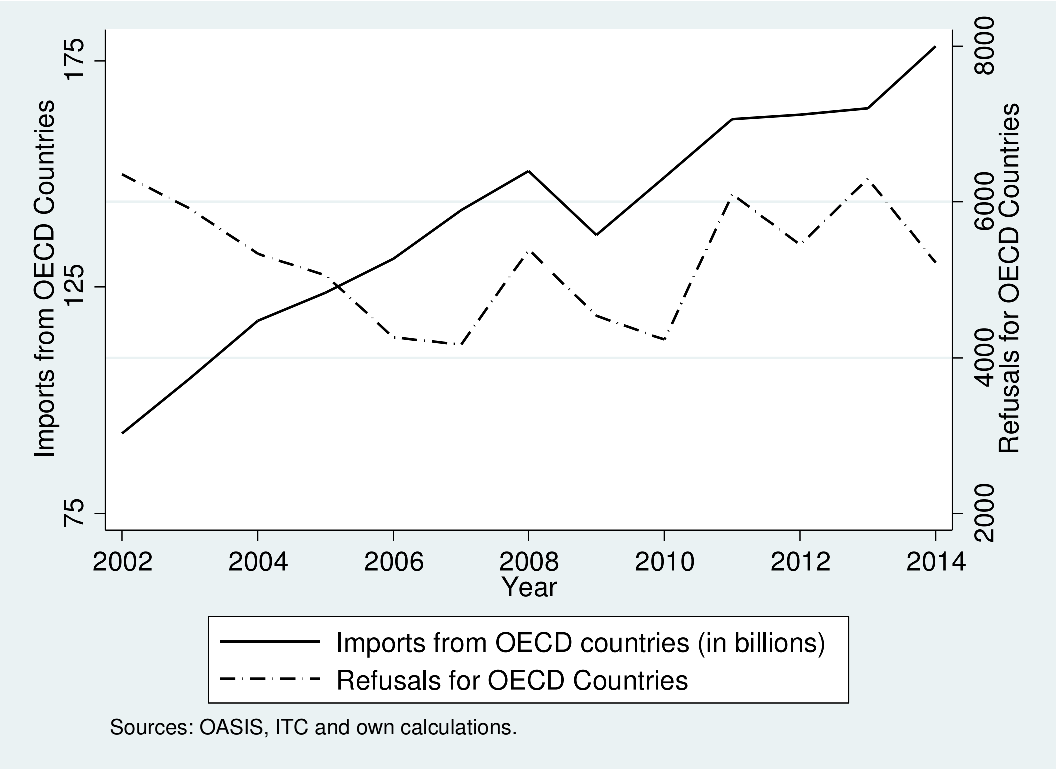
Graph 1: U.S. Unemployment Rate and Hidden Protectionism?



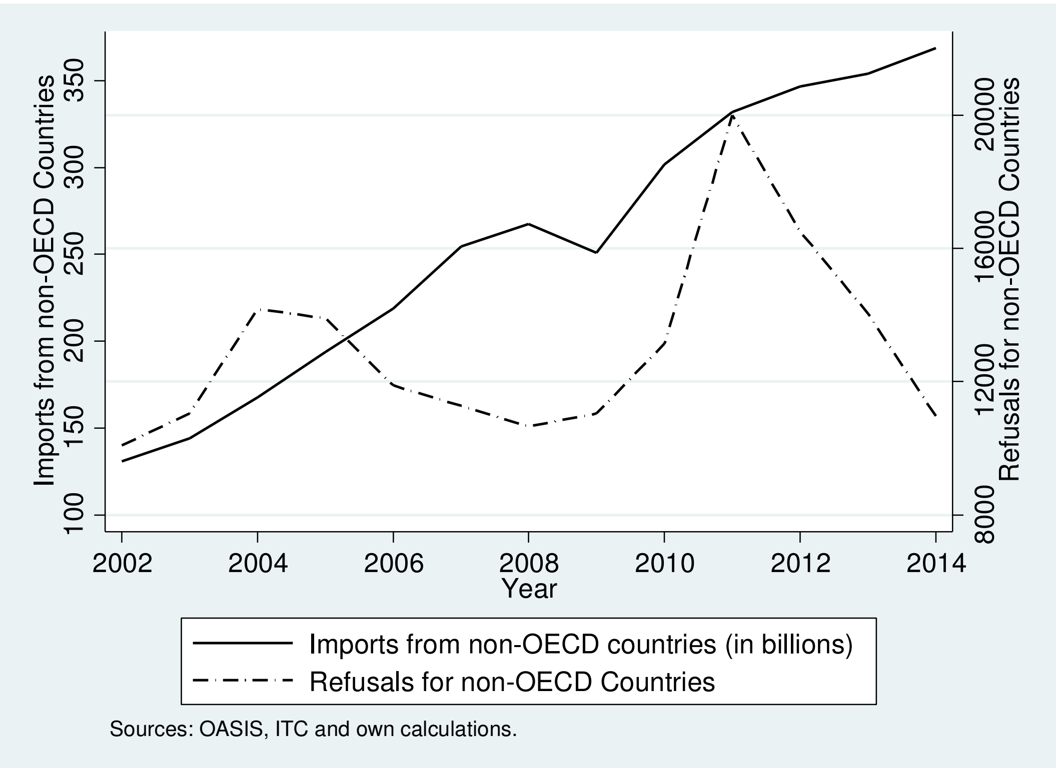
Graph 2: Total U.S. Imports and Refusals of FDA Regulated Products



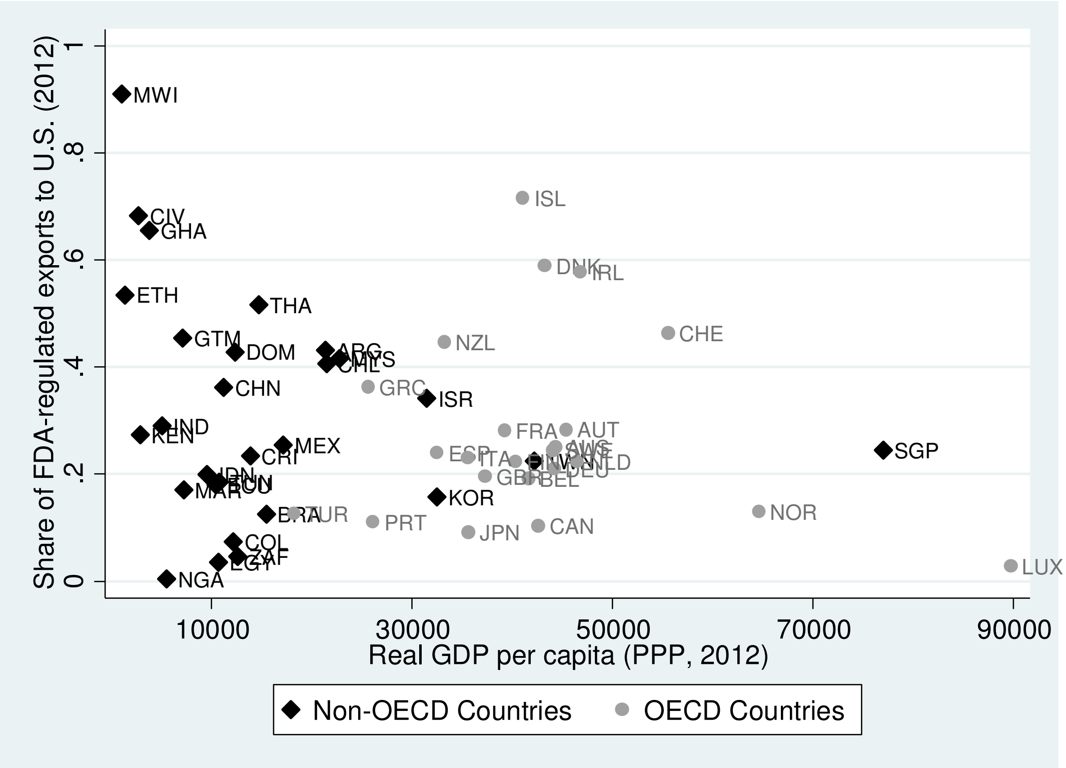
Graph 3: U.S. Imports and Refusals for OECD Countries



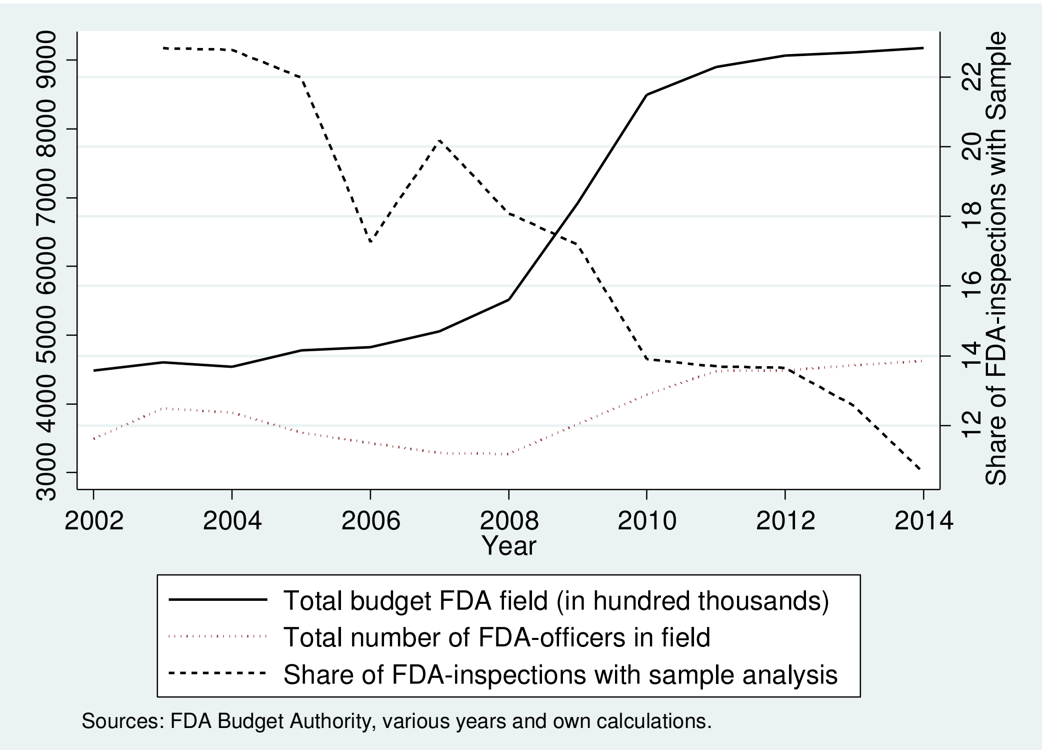
Graph 4: U.S. Imports and Refusals for non-OECD Countries



Graph 5: Share of FDA-regulated Products in Total Exports to United States (2012)

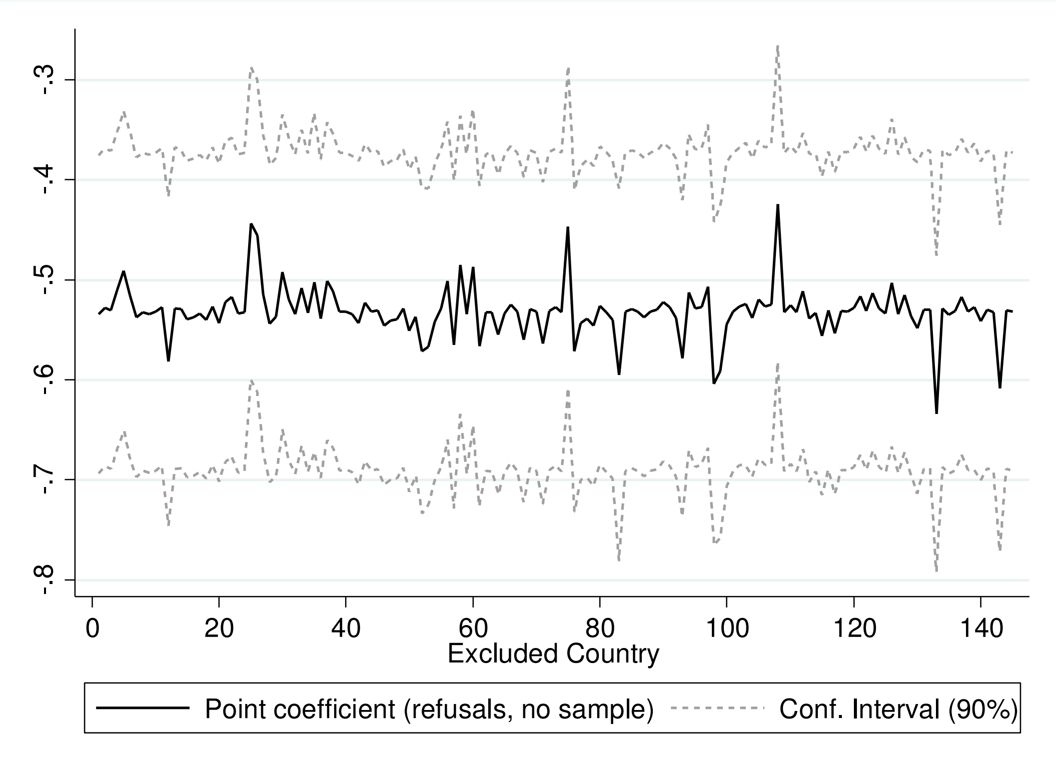


Graph 6: FDA-budget for Field Activities, FDA-officers in the Field and FDA-inspections with Product Sample Analysis

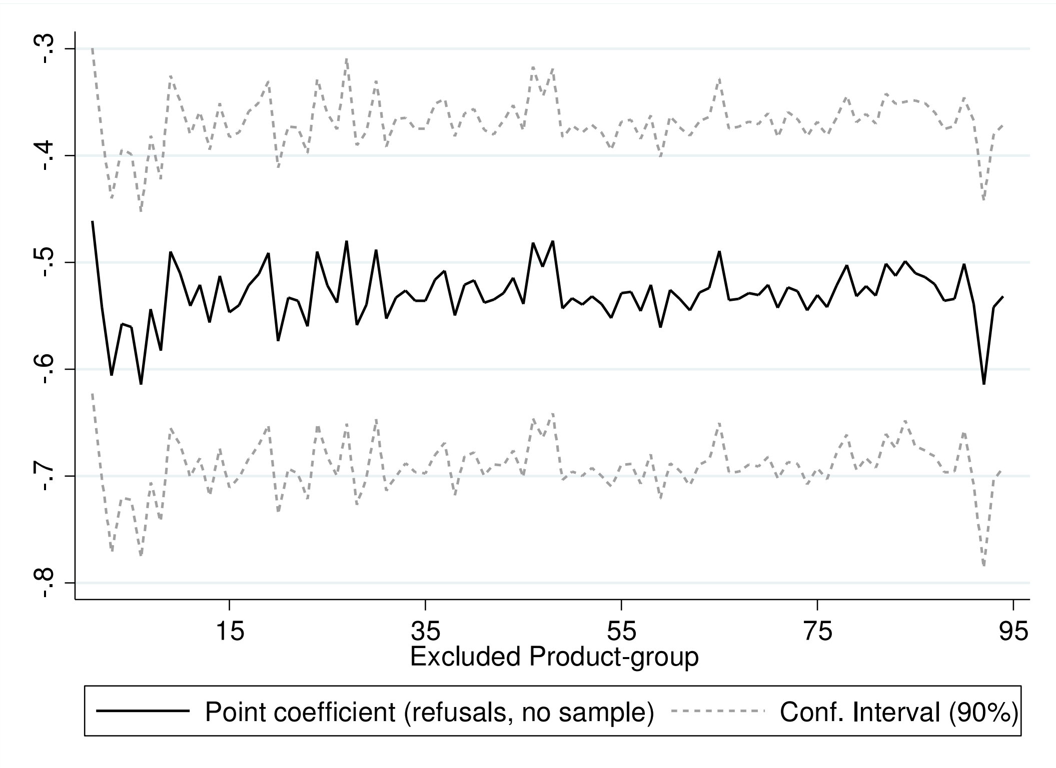


Sources: FDA Budget Authority, various years and own calculations.

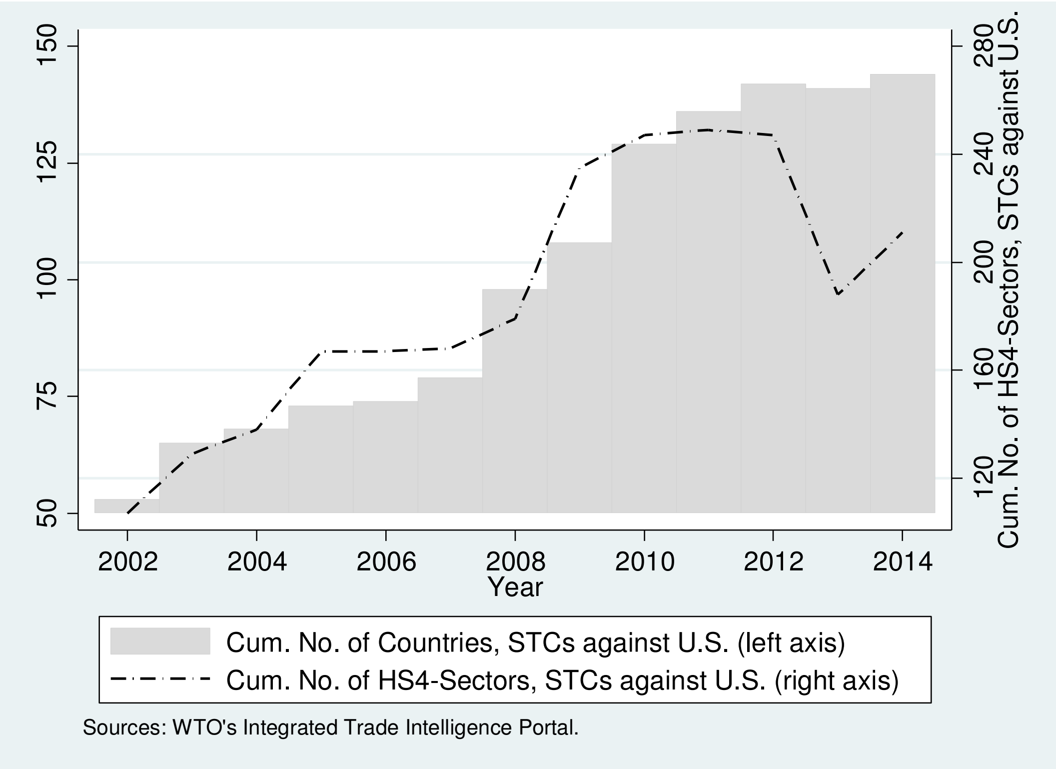
Graph 7: U.S. Imports and Refusals (without sample analysis) – Robustness (Countries), Non-OECD Countries (Crisis Years, 2008-2012)



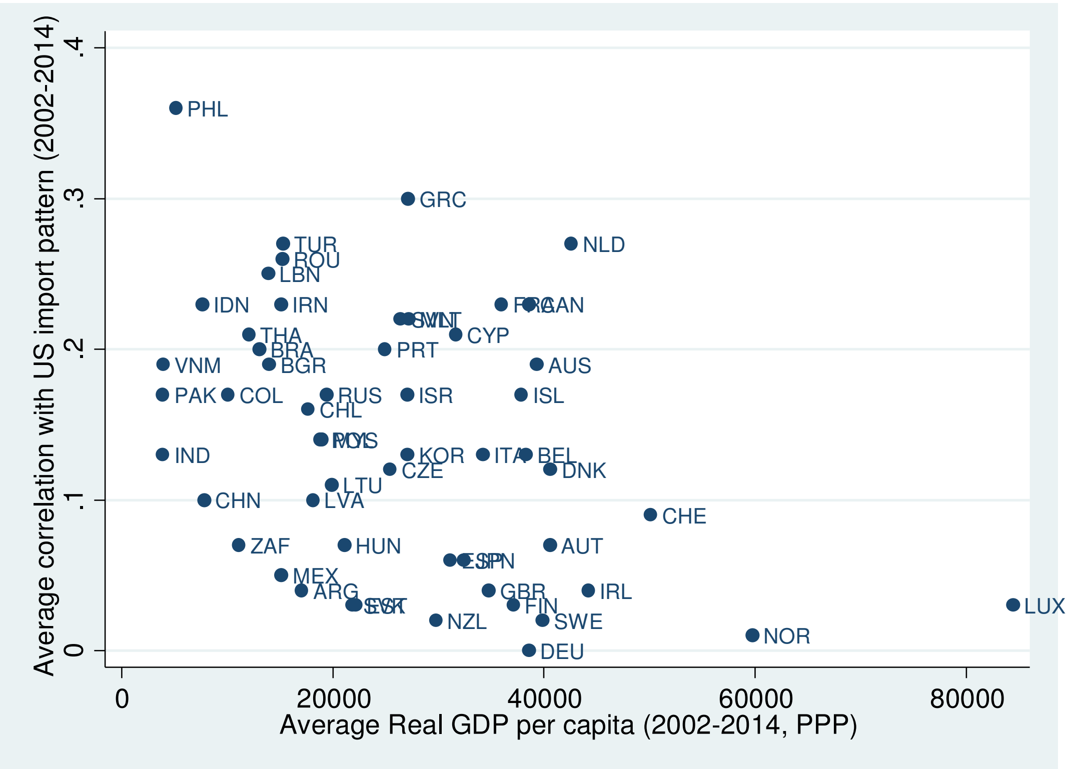
Graph 8: U.S. Imports and Refusals (without sample analysis) – Robustness (Product-groups), Non-OECD Countries (Crisis Years, 2008-2012)



Graph 9: Cumulative Number of Specific Trade Concerns (STCs) against the United States



Graph 10: Correlations between United States and Other High Income Countries' Import Patterns (average 2002-2014)



Appendix

Table A1: Number of U.S. Import Refusals per Matched Product-group 2002 – 2014 (Top Ten out of 93 Product-groups)

Product-group Number	*Description*	*Number of Import Refusals*
84	Other medicaments, except antibiotics and hormones	38610
90	Medical instruments, machines and other medical devices	35381
74	Skin care and make up	10135
36	Bread and pastry, pudding, other baker ware	9423
2	Fish, dried, salted, smoked or in brine	8488
46	Sugar confectionary without cacao	7705
3	Crustaceans, fresh, chilled, frozen, dried, smoked or in brine	7233
60	Sauces, mixed dressings and condiments	6883
8	Fruits used as vegetables, fresh or chilled	6241
1	Fish, fresh, chilled or frozen	6121

Table A2: Reasons of U.S. Import Refusals (Top Ten Reasons from 2002 - 2014)

FDA Reason Code	*FDA Reason Description*	*Number of Import Refusals*
NOT LISTED	It appears the drug or device is not included in a list required by Section 510(j), or a notice or other information respecting it was not provided as required by section 510(j) or 510(k).	41596
UNAPPROVED	The article appears to be a new drug without an approved new drug application.	37323
FILTHY	The article appears to consist in whole or in part of a filthy, putrid, or decomposed substance or be otherwise unfit for food.	24499
NUTRIT LBL	The article appears to be misbranded in that the label or labeling fails to bear the required nutrition information.	21364
NO PROCESS	It appears that the manufacturer has not filed information on its scheduled process as required by 21 CFR 108.25(c)(2) or 108.35(c)(2).	17334
UNSAFE COL	The article appears to be a color additive for the purposes of coloring only in or on drugs or devices, and is unsafe within the meaning of Section 721(a).	15960
SALMONELLA	The article appears to contain Salmonella, a poisonous and deleterious substance which may render it injurious to health.	15598
LIST INGRE	It appears the food is fabricated from two or more ingredients and the label does not list the common or usual name of each ingredient.	12939
REGISTERED	It appears the device is subject to listing under 510(j) and the initial distributor has not registered as required by 21 CFR 807.20 (a)(5).	12412
NEEDS FCE	It appears the manufacturer is not registered as a low acid canned food or acidified food manufacturer pursuant to 21 CFR 108.25(c)(1) or 108.35(c)(1).	11603

Table A3: U.S. Imports and Refusals Dummy – Different Estimators

Variables	***OLS***	***Fixed effects***	***Arellano-Bond***	***Arellano-Bond***
	(1)	***(2)***	***(3)***	***(4)***
Log Imports (t-1)	0.731**	0.449**	0.416**	0.421**
	(0.01)	(0.01)	(0.03)	(0.03)
Log Imports (t-2)	0.241**	0.061**	0.058**	0.062**
	(0.01)	(0.01)	(0.02)	(0.02)
Dummy refusal (t-1)	0.045**	0.028*	-0.002	-0.551°
	(0.01)	(0.01)	(0.01)	(0.31)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	No	Yes	Yes	Yes
Refusals endogenous	No	No	No	Yes
AR(1) (*p*-value)			0.000	0.000
AR(2) (*p*-value)			0.711	0.754
Sargan-test (*p*-value)			0.600	0.726
Hansen-test (*p*-value)			0.843	0.797
Number of instruments			15	17
Number of groups			3468	3468
Number of countries	170	170	167	167
Number of observations	33787	33787	29882	29882

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable dummy refusal takes the value of one, if in a given product-group k from country i at least one refusal incidence is recorded in year t. Robust standard errors are reported in parentheses. The estimates in Column (1) and (2) are based on pooled OLS and (country-product) fixed effects, with standard errors being clustered at the country-level. Columns (3) and (4) employ a two-step Arellano-Bond estimator with robust standard errors. The lagged dependent variable is instrumented with its first through third lag. In Column (3), we define dummy refusal as exogenous. In Column (4), we allow the dummy refusal to be endogenous and instrument it with its second through fourth lag. The instrument matrix is collapsed. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table A4: U.S. Imports and Refusals – Further Fixed Effects

Variables	***Baseline***	***WTO notifications***	***Further Fixed Effects I***	***Further Fixed Effects II***
	(1)	***(2)***	***(3)***	***(4)***
Log Imports (t-1)	0.419**	0.418**	0.382**	0.310**
	(0.03)	(0.03)	(0.04)	(0.05)
Log Imports (t-2)	0.059**	0.060**	0.044**	0.027
	(0.01)	(0.02)	(0.02)	(0.02)
Refusals (t-1)	-0.134°	-0.135°	-0.171*	-0.212°
	(0.08)	(0.08)	(0.09)	(0.11)
WTO notifications		-0.034*	-0.049*	-0.046*
		(0.02)	(0.02)	(0.02)
WTO notifications (t-1)		-0.012	0.016	0.011
		(0.02)	(0.02)	(0.02)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
Sector-year FE	No	No	Yes	Yes
Country-time FE	No	No	No	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.669	0.669	0.402	0.316
Sargan-test (*p*-value)	0.875	0.865	0.313	0.346
Hansen-test (*p*-value)	0.950	0.951	0.620	0.587
Number of instruments	17	19	59	394
Number of groups	3468	3468	2937	2937
Number of countries	167	167	68	68
Number of observations	29882	29882	26032	26032

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. In Column (1), we report the baseline results from Table 1, Column (4), for comparability. Column (2) adds the variable WTO notifications, which is a dummy variable taking the value of one if a given product-group is affected by at least one WTO notification (TBT or SPS agreement) of the United States in year t and zero otherwise. These variables are assumed to be exogenous. The specifications presented in Columns (3) and (4) allow for further fixed effects. Since these specifications require more cross-sectional variation within countries, we restrict the sample to countries with export flows to the United States in at least 20 out of the 93 product-groups. Column (3) extends the baseline with sector-year fixed effects, where all product-groups are classified into five more aggregated sectors. Column (4) includes sector-year fixed effects and additionally country-time fixed effects, with time being defined as two-year spells. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table A5: U.S. Imports and Refusals – Further Specification Tests

	Baseline	***External Instrument I***	***External Instrument II***	***External Instrument III***
Upper Panel (Arellano-Bond)	***(1)***	***(2)***	***(3)***	***(4)***
Refusals (t-1)	-0.134°	-0.134°	-0.135°	-0.136°
	(0.08)	(0.08)	(0.08)	(0.08)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.669	0.668	0.668	0.682
Sargan-test (*p*-value)	0.875	0.943	0.941	0.929
Hansen-test (*p*-value)	0.950	0.982	0.980	0.967
Number of instruments	17	18	18	18
Number of groups	3468	3468	3468	3468
Number of countries	167	167	167	167
Number of observations	29882	29882	29882	29882
Lower Panel (2 SLS)	***(5)***	***(6)***	***(7)***	***(8)***
Refusals (t-1)	-0.172	-0.172	-0.165	-0.159
	(0.11)	(0.11)	(0.11)	(0.11)
Hansen-test (*p*-value)	0.768	0.758	0.689	0.597
Number of instruments	15	16	16	16
Number of groups	3197	3197	3197	3197
Number of countries	160	160	160	160
Number of observations	22950	22950	22950	22950
Under-identification test (rk LM stat)	83.29	83.38	84.03	84.32
Under-identification test (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
Weak identification test (Wald rk F stat)	14.09	12.08	12.16	12.20
H_0: relative OLS bias>30% (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW Chi-sq test (imports) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW F test (imports) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW Chi-sq test (refusals) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW F test (refusals) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations in the **upper panel** are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable (included in the regression, but not reported in the table) is instrumented with the first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. The baseline results from Table 1, Column (4) are reported in Column (1). Columns (2), (3) and (4) report the results for the baseline specification, where we additionally use all information (2) and warnings and refusals (3) from the RASFF-member states Finland, Germany, Norway and Sweden as well as warnings and refusals (4) from Germany as external instruments, entering the specification with the first lag. All estimates in the **lower panel** are based on a two-stage least squares (2SLS) estimator, whereby we replicate the GMM-estimator in this IV-framework. We report the results for our main variable of interest (refusals), the Hansen-test of over-identifying restrictions as well as a number of further specification tests on under-idenfication and weak instruments. The null hypothesis of the Kleibergen-Paap LM test is that the structural equation is under-identified (see Kleibergen and Paap, 2006). We also report the Kleibergen-Paap Wald statistics for the weak instruments tests, where the null hypothesis is that instruments are (jointly) weak (see Kleibergen and Paap, 2006). We follow Bazzi and Clemens (2013) and report the p-values for the null hypotheses of these two tests that the bias in the point estimates of the endogenous variables is greater than 30 percent of the OLS bias. Finally, the Sanderson-Windmeijer (SW) chi-squared and F statistics test the null hypotheses of under-identification and weak instruments for each endogenous regressor (here imports and refusals) separately (see Sanderson and Windmeijer, 2015). **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table A6: U.S. Imports and Refusals – Further Fixed Effects, Non-OECD Countries

Variables	*Baseline*	*WTO notifications*	*Further Fixed Effects I*	*Further Fixed Effects II*
	(1)	*(2)*	*(3)*	*(4)*
Log Imports (t-1)	0.386**	0.382**	0.337**	0.302**
	(0.04)	(0.04)	(0.05)	(0.05)
Log Imports (t-2)	0.059**	0.059**	0.042*	0.033
	(0.02)	(0.02)	(0.02)	(0.02)
Refusals (t-1)	-0.204*	-0.210*	-0.240*	-0.243*
	(0.10)	(0.10)	(0.10)	(0.12)
WTO notifications		-0.050*	-0.077**	-0.076*
		(0.02)	(0.03)	(0.03)
WTO notifications (t-1)		-0.039*	-0.023	-0.026
		(0.02)	(0.03)	(0.03)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
Sector-year FE	No	No	Yes	Yes
Country-time FE	No	No	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.393	0.405	0.174	0.145
Sargan-test (*p*-value)	0.936	0.933	0.066	0.084
Hansen-test (*p*-value)	0.930	0.932	0.268	0.281
Number of instruments	17	19	59	294
Number of groups	2496	2496	1992	1992
Number of countries	144	144	48	48
Number of observations	20839	20839	17216	17216

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. In Column (1), we report the baseline results from Table 2, Column (4), for comparability. Column (2) adds the variable WTO notifications, which is a dummy variable taking the value of one if a given product-group is affected by at least one WTO notification (TBT or SPS agreement) of the United States in year t and zero otherwise. These variables are assumed to be exogenous. The specifications presented in Columns (3) and (4) allow for further fixed effects. Since these specifications require more cross-sectional variation within countries, we restrict the sample to countries with export flows to the United States in at least 20 out of the 93 product-groups. Column (3) extends the baseline with sector-year fixed effects, where all product-groups are classified into five more aggregated sectors. Column (4) includes sector-year fixed effects and additionally country-time fixed effects, with time being defined as two-year spells. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table A7: U.S. Imports and Refusals – Further Specification Tests, Non-OECD Countries

	Baseline	*External Instrument I*	*External Instrument II*	*External Instrument III*
Upper Panel (Arellano-Bond)	***(1)***	***(2)***	***(3)***	***(4)***
Refusals (t-1)	-0.204*	-0.203*	-0.208*	-0.209*
	(0.10)	(0.10)	(0.10)	(0.10)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.393	0.394	0.392	0.399
Sargan-test (*p*-value)	0.936	0.974	0.972	0.969
Hansen-test (*p*-value)	0.930	0.973	0.966	0.961
Number of instruments	17	18	18	18
Number of groups	2496	2496	2496	2496
Number of countries	144	144	144	144
Number of observations	20839	20839	20839	20839
Lower Panel (2 SLS)	***(5)***	***(6)***	***(7)***	***(8)***
Refusals (t-1)	-0.231°	-0.230°	-0.229°	-0.232°
	(0.13)	(0.13)	(0.13)	(0.13)
Hansen-test (*p*-value)	0.753	0.861	0.757	0.789
Number of instruments	15	16	16	16
Number of groups	2262	2262	2262	2262
Number of countries	137	137	137	137
Number of observations	15855	15855	15855	15855
Under-identification test (rk LM stat)	81.70	81.70	81.88	81.73
Under-identification test (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
Weak identification test (Wald rk F stat)	14.16	12.15	12.14	12.15
H_0: relative OLS bias>30% (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW Chi-sq test (imports) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW F test (imports) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW Chi-sq test (refusals) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW F test (refusals) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations in the **upper panel** are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable (included in the regression, but not reported in the table) is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. The baseline results from Table 1, Column (4) are reported in Column (1). Columns (2), (3) and (4) report the results for the baseline specification, where we additionally use all information (2) and warnings and refusals (3) from the RASFF-member states Finland, Germany, Norway and Sweden as well as warnings and refusals (4) from Germany as external instruments, entering the specification with the first lag. All estimates in the **lower panel** are based on a two-stage least squares (2SLS) estimator, whereby we replicate the GMM-estimator in this IV-framework. We report the results for our main variable of interest (refusals), the Hansen-test of over-identifying restrictions as well as a number of further specification tests on under-idenfication and weak instruments. The null hypothesis of the Kleibergen-Paap LM test is that the structural equation is under-identified (see Kleibergen and Paap, 2006). We also report the Kleibergen-Paap Wald statistics for the weak instruments tests, where the null hypothesis is that instruments are jointly weak (see Kleibergen and Paap, 2006). We follow Bazzi and Clemens (2013) and report the p-values for the null hypotheses of these two tests that the bias in the point estimates of the endogenous variables is greater than 30 percent of the OLS bias. Finally, the Sanderson-Windmeijer (SW) chi-squared and F statistics test the null hypotheses of under-identification and weak instruments for each endogenous regressor (here imports and refusals) separately (see Sanderson and Windmeijer, 2015). **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table A8: U.S. Imports and Refusals – Further Fixed Effects, Non-OECD Countries (Crisis Years, 2008-2012)

Variables	***Baseline***	***WTO notifications***	***Further Fixed Effects I***	***Further Fixed Effects II***
	(1)	***(2)***	***(3)***	***(4)***
Log Imports (t-1)	0.385**	0.387**	0.344**	0.340**
	(0.07)	(0.07)	(0.09)	(0.10)
Log Imports (t-2)	0.067*	0.067*	0.051	0.047
	(0.03)	(0.03)	(0.03)	(0.03)
Refusals (Sample) (t-1)	0.326	0.329	0.313	0.318
	(0.26)	(0.26)	(0.28)	(0.26)
Refusals (no Sample) (t-1)	-0.532*	-0.525*	-0.533*	-0.505°
	(0.26)	(0.25)	(0.24)	(0.26)
WTO notifications		0.003	-0.027	-0.029
		(0.03)	(0.04)	(0.04)
WTO notifications (t-1)		0.018	0.033	0.027
		(0.03)	(0.04)	(0.03)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
Sector-year FE	No	No	Yes	Yes
Country-time FE	No	No	No	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.311	0.314	0.391	0.395
Sargan-test (*p*-value)	0.911	0.912	0.456	0.446
Hansen-test (*p*-value)	0.931	0.932	0.773	0.754
Number of instruments	15	17	37	129
Number of groups	2334	2334	1883	1883
Number of countries	144	144	48	48
Number of observations	10491	10491	8640	8640

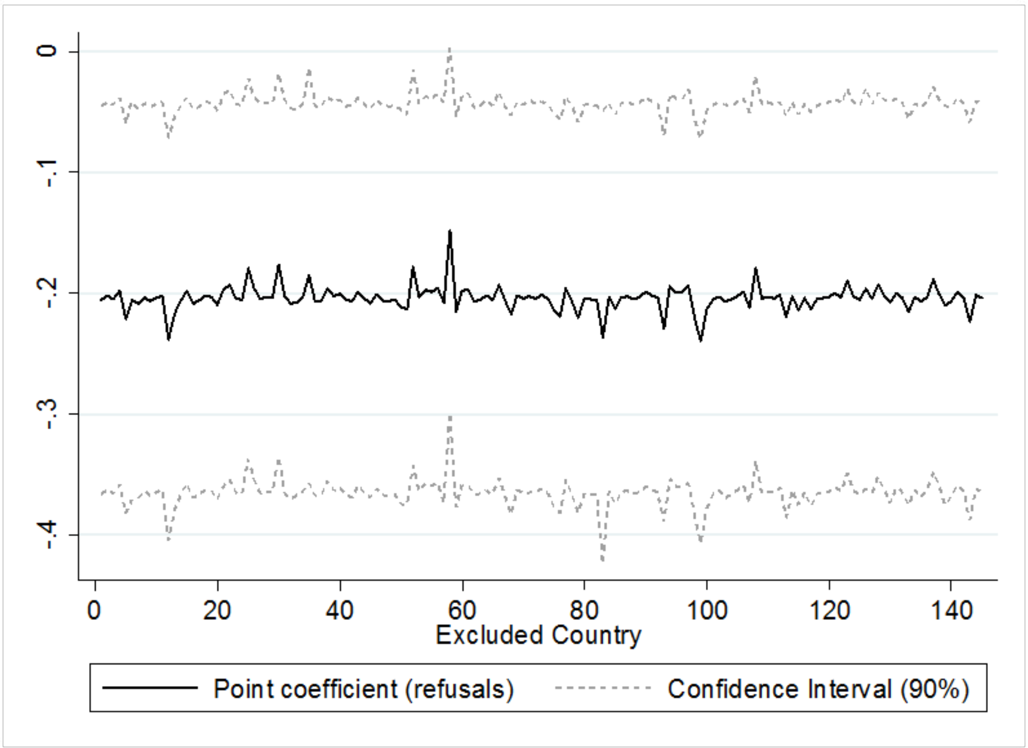
Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). All estimations are based on the two-step Arellano-Bond estimator with robust standard errors reported in parentheses. The lagged dependent variable is instrumented with its first through third lag. The variable refusals is allowed to be endogenous and is instrumented with its second through fourth lag. The instrument matrix is collapsed. In Column (1), we report the baseline results from Table 5, Column (4), for comparability. Column (2) adds the variable WTO notifications, which is a dummy variable taking the value of one if a given product-group is affected by at least one WTO notification (TBT or SPS agreement) of the United States in year t and zero otherwise. These variables are assumed to be exogenous. The specifications presented in Columns (3) and (4) allow for further fixed effects. Since these specifications require more cross-sectional variation within countries, we restrict the sample to countries with export flows to the United States in at least 20 out of the 93 product-groups. Column (3) extends the baseline with sector-year fixed effects, where all product-groups are classified into five more aggregated sectors. Column (4) includes sector-year fixed effects and additionally country-time fixed effects, with time being defined as two-year spells. **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Table A9: U.S. Imports and Refusals – Further Specification Tests, Non-OECD Countries (Crisis Years, 2008-2012)

	Baseline	*External Instrument I*	*External Instrument II*	*External Instrument III*
Upper Panel (Arellano-Bond)	***(1)***	***(2)***	***(3)***	***(4)***
Refusals (sample) (t-1)	0.326	0.300	0.326	0.335
	(0.26)	(0.25)	(0.26)	(0.26)
Refusals (no sample) (t-1)	-0.532*	-0.501*	-0.532*	-0.541*
	(0.26)	(0.25)	(0.25)	(0.26)
Time FE	Yes	Yes	Yes	Yes
Country-Product FE	Yes	Yes	Yes	Yes
Refusals endogenous	Yes	Yes	Yes	Yes
AR(1) (*p*-value)	0.000	0.000	0.000	0.000
AR(2) (*p*-value)	0.311	0.347	0.313	0.304
Sargan-test (*p*-value)	0.911	0.892	0.954	0.941
Hansen-test (*p*-value)	0.931	0.917	0.967	0.953
Number of instruments	15	16	16	16
Number of groups	2334	2334	2334	2334
Number of countries	144	144	144	144
Number of observations	10491	10491	10491	10491
Lower Panel (2 SLS)	***(5)***	***(6)***	***(7)***	***(8)***
Refusals (sample) (t-1)	0.240	0.196	0.235	0.246
	(0.30)	(0.29)	(0.29)	(0.30)
Refusals (no sample) (t-1)	-0.483°	-0.458°	-0.483°	-0.485°
	(0.25)	(0.25)	(0.25)	(0.25)
Hansen-test (*p*-value)	0.866	0.866	0.922	0.925
Number of instruments	15	16	16	16
Number of groups	2148	2148	2148	2148
Number of countries	137	137	137	137
Number of observations	9903	9903	9903	9903
Under-identification test (rk LM stat)	21.79	21.64	21.75	21.66
Under-identification test (*p*-value)	0.003	0.006	0.005	0.006
Weak identification test (Wald rk F stat)	2.54	2.28	2.82	2.82
H_0: relative OLS bias>30% (*p*-value)	0.502	0.638	0.637	0.639
SW Chi-sq test (imports) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW F test (imports) (*p*-value)	< 0.001	< 0.001	< 0.001	< 0.001
SW Chi-sq test (refusals) (*p*-value)	0.003	0.005	0.005	0.005
SW F test (refusals) (*p*-value)	0.003	0.005	0.005	0.006

Notes: Dependent variable: Log imports of 93 product-groups to the United States from 2002-2014. The variable refusals refers to the total number of refusals in a given product-group k from country i in year t and enters as the log(1+refusals). For further details on the **upper panel** are reported in the Table A5 or A7. All estimates in the **lower panel** are based on a two-stage least squares (2SLS) estimators, whereby we replicate the GMM-estimator in this IV-framework. We report the results for our main variable of interest (refusals), the Hansen-test of over-identifying restrictions as well as a number of further specification tests on under-idenfication and weak instruments. The null hypothesis of the Kleibergen-Paap LM test is that the structural equation is under-identified (see Kleibergen and Paap, 2006). We also report the Kleibergen-Paap Wald statistics for the weak instruments tests, where the null hypothesis is that instruments are jointly weak (see Kleibergen and Paap, 2006). We follow Bazzi and Clemens (2013) and report the p-values for the null hypotheses of these two tests that the bias in the point estimates of the endogenous variables is greater than 30 percent of the OLS bias. Finally, the Sanderson-Windmeijer (SW) chi-squared and F statistics test the null hypotheses of under-identification and weak instruments for each endogenous regressor, whereby we report here only the higher p-value of the two refusals variables (see Sanderson and Windmeijer, 2015). **, * and ° denotes significant at the 1%, 5% and 10% level, respectively.

Graph A1: U.S. Imports and Refusals – Robustness (Countries), Non-OECD Countries



Graph A2: U.S. Imports and Refusals – Robustness (Product-groups), Non-OECD Countries