Consequences of FTA Withdrawal: Evidence from "Uxit"

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Abstract

We use a unique case study to estimate the effect of withdrawing from a free trade agreement on international trade. Lately, the political opposition to international economic cooperation has been on the rise, but little is known about how the withdrawal from a trade agreement affects trade. We analyze a quasi-natural experiment to provide first empirical evidence. In 2004, Estonia joined the European Union, which mandated that it withdraws from its FTA with Ukraine ("Uxit"). Based on the gravity model of international trade, we provide evidence from triple difference–in–differences as well as PPML panel estimations that trade volumes between Estonia and Ukraine fell by more than 20%. We find that withdrawing an FTA revokes all benefits and that no institutional memory is left behind. General equilibrium estimates suggest that FTA withdrawal led to a noticeable loss in welfare of members.

Keywords: free trade agreement, withdrawal, gravity, European Union, Estonia, Ukraine
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1. Introduction

Until recently, international economic integration via Free Trade Agreements (FTAs) largely has been a one–way street. Almost at a yearly basis, new FTAs were signed and trade barriers between countries were lowered. At the same time, the literature kept providing ever new evidence for the positive effects of FTAs. Most prominently, FTAs have been shown to increase the trade volumes between countries (Baier and Bergstrand, 2007; Baier, Yotov, and Zylkin, 2019; Eicher, Henn, and Papageorgiou, 2012), raise the quality and variety of products available to consumers (Berlingieri et al., 2018; Broda and Weinstein, 2006), spur stock markets (Moser and Rose, 2014), and increase countries' overall welfare and economic efficiency (Anderson and Yotov, 2016; Khandelwal et al., 2013). We do not know yet, however, how the withdrawal of an FTA affects its former signatories. Are all the FTA effects undone upon withdrawal, or is some institutional memory left behind?

Currently, trade agreements are facing severe political backlashes across the world. Protests against FTAs abound, culminating in a new wave of economic nationalism, which associates economic integration with rent-seeking behavior and suspects re-distributional side effects that hurt the already less well-off (Rodrik, 2018). As a consequence, both existing and planned FTAs undergo reconsideration. Among other things, the negotiations of large-scale FTAs like the Trans-Atlantic Trade and Investment Partnership (TTIP) between the European Union and the United States have been halted. Further, the United States withdrew from the North American Free Trade Agreement (NAFTA) with Canada and Mexico, only to renegotiate a "better deal", which is now called the United States-Mexico-Canada Agreement (USMCA). Additionally, the USA engaged in a trade war with China and Europe, accepting a sizeable reduction in consumer welfare in return for gains for domestic producers (Fajgelbaum et al., 2020). In a similar vein, the United Kingdom officially withdrew from the European Union in January 2020 and left the EU Single Market and Customs Union in January 2021. In response, economic research starts shifting away from the focus on the *formation* of FTAs and towards understanding the economic effects of their *dissolution*.

In this paper, we exploit a quasi-natural experiment to estimate the effect of FTA withdrawal on trade and welfare. With the Eastern Enlargement of the European Union in 2004,[1], the new member states joined the European Customs Union with its centralized competence for trade agreements; in other words, member states replaced their earlier national trade agreements with those of the Union on the day of accession. This caused little upset to existing trade agreements by the new members, since their trade agreements had been negotiated with EU accession in mind.[2] However, there is one *single* exception: Estonia, one of the new

[1]See Gateva (2016, ch. 2) for a discussion of the "A10" Enlargement Process.

[2]For example, the "A10" states had an FTA among each other; upon accession, these agreements were "upgraded" to the EU Common Market. The A10 also had an FTA with the European Free Trade Association (EFTA); after EU accession, the countries continued to enjoy an FTA with the EFTA countries – since the EU also had an FTA with them in place.

EU members, had an FTA in force with Ukraine since 1999. Because the EU did not have an FTA with Ukraine at the time Estonia joined the Union, Estonia had to withdraw from this agreement as part of its EU accession process. This "Uxit" was – as we argue below – driven by considerations exogenous to Ukraine–Estonian bilateral characteristics. We provide results from a triple difference–in–differences estimator as well as PPML regressions to identify the causal impact of FTA withdrawal on trade and welfare.

We rely on two different estimation methods and samples in the analysis. Leveraging high-frequency monthly data, we estimate triple difference-in-differences tetrad regressions and find that the withdrawal of the FTA between Estonia and Ukraine caused a loss of 20%–24% in bilateral trade. Additionally, we provide PPML estimates based on yearly trade data which suggest an even larger withdrawal-effect of up to -49%. This effect is especially predominant in the transport sector, the main trading-sector between Estonia and Ukraine. Comparing our results to Baier, Yotov, and Zylkin (2019), who find an average FTA trade creation effect of 26%, we interpret our finding as evidence that essentially *all* FTA trade gains become undone after FTA withdrawal and that no 'institutional memory' outlasts the agreement, as was hypothesized in Head et al. (2010). Further, we follow Baier, Yotov, and Zylkin (2019) and estimate directional effects of the FTA withdrawal, in effect disentangling the trade-reduction effect on Estonian exports from the effect on Ukrainian exports. We find no significant differences with respect to the direction of trade, with exporters from both countries reducing their shipments to the former FTA-partner country by around 40%. While there is clear evidence that trade started to decrease already shortly prior to the FTA withdrawal, suggesting an anticipation effect, we find no evidence of a delay in trade reduction afterwards. A Chow test fails to find evidence for a gradual fall in trade. Rather, the effect appears directly after the withdrawal and remains constant over time.

Finally, general equilibrium calculations suggest that both countries faced severe welfare losses due to the FTA withdrawal, whereas Estonia's Baltic trade partners enjoyed noticeable welfare increases. Incorporating the simultaneous EU accession effect into our welfare analysis, we find that Estonia enjoyed a net welfare gain from the EU accession, despite lower than its neighbor countries, whereas Ukraine suffered an overall welfare loss of -0.16%.

We contribute to a new but growing literature that aims to quantify the effects of economic disintegration. Especially the UK's withdrawal from the EU has spurred academic interest to understand the economic consequences of re-establishing barriers to trade. Already since they joined the European Union, the UK's public discontent with the EU drew scholars' interest (see, e.g., Pain and Young, 2004). The scientific attention peaks since 2016 after the UK eventually voted to leave the EU. Numerous recent papers focus on the imminent Brexit to quantify the trade- and welfare effects of dissolving a trade agreement. However, the fact that Brexit did not actually happen until 2020 and its full effects will unfold from 2021 onward complicates this task. The most common approach is to first estimate the effect of the UK's EU membership on trade and welfare, and then construct a counterfactual world in which the

UK was not part of the EU. Then, by comparing the actual UK to its simulated counterfactual, researchers derive an estimate for the effect of the UK's withdrawal on its trade with the EU as well as its overall welfare. These papers find a potential trade reduction between the UK and the EU in the range of 25% and 45% (Oberhofer and Pfaffermayr, 2017) as well as welfare-losses for the British economy between 1% and 9.4% (Dhingra et al. 2017, Felbermayr, Gröschl, and Steininger 2018), depending on the assumptions made and models used. Loosening the focus on the UK but staying in Europe, other papers used similar techniques to construct a counterfactual Europe where the European Union with its Single Market, Currency Union and open borders did not exist. These papers show that European countries would lose up to 23% of overall welfare (Felbermayr, Gröschl, and Heiland, 2018; Mayer, Vicard, and Zignago, 2019). Leaving Europe entirely, Baier, Bergstrand, and Bruno (2019) estimate a potential welfare loss of up to 2.1% for the Canadian economy from the dissolution of NAFTA, had it not been replaced by the new and "better" USMCA agreement. It therefore seems fair to conclude that removing – or never signing – a free trade agreement has large negative effects on the involved economies. However, all results reported so far rely on simulations or informed guesses to derive their conclusions, since there is not (yet) a counterfactual world where e.g. the UK left the common market. Hence, with this paper we close this gap by estimating the trade effects of the withdrawal from an FTA using a quasi–experiment based on a real world example.

Our findings are further important in the context of future UK trade policy after Brexit. Existing analyses suggest that the effects of Brexit critically depend on the terms of trade established post withdrawal (Dhingra et al. 2017, Felbermayr, Gröschl, and Steininger 2018). Even though the EU-UK Trade and Cooperation Agreement struck last-minute in December 2020 stipulates zero tariffs and free entry of goods, it is still much less comprehensive than the Single Market and Customs Union.[3] Our evidence from Estonia suggests that even when MFN tariffs are close to zero, a comprehensive FTA can give a substantial boost to trade. Prior to EU accession, Estonia was close to pursuing unilateral free trade;[4] however, the Estonia-Ukraine FTA did grant tariff-free access for all tariff lines. This suggests that particularly comprehensive agreements may have extra trade-promoting effects.

This paper proceeds by describing the Estonia–Ukraine FTA in Section 2. Our estimation strategy is outlined in the following Section 3, while Section 4 turns to results. In Section 5, we present a couple of extensions and robustness tests. Finally, Section 6 concludes.

[3] To quote the European Commission directly: "The EU and the UK will form two separate markets; two distinct regulatory and legal spaces. This will create barriers to trade in goods and services and to cross-border mobility and exchanges that do not exist today – in both directions." (European Commission Press release December 24, 2020, https://ec.europa.eu/commission/presscorner/detail/en/IP_20_2531).

[4] Mean MFN tariffs were very low below 2%. See also Feldmann and Sally (2002) for an in-depth analysis.

2. The Estonia–Ukraine Free Trade Agreement

Estonia is a small open economy bordered by the Baltic Sea, Russia and Latvia. After the fall of the Soviet Union, Estonia transitioned rapidly to a market economy. Today, it is considered one of the most successful post-socialist economies (Norkus, 2007). Prior to joining the European Union in 2004, Estonia practiced a very liberal trade policy: according to the *World Trade Organization*, its average MFN tariff was only 1.68% in 2002 – and for 93% of tariff lines, Estonia granted tariff–free access on an MFN basis. In other words, Estonian tariffs were unusually low by international standards. Additionally, Estonia had free trade agreements in place with the European Union, the EFTA countries and Ukraine. These agreements were unusually comprehensive, since Estonia granted tariff–free access on *all* goods to *each* FTA partner.

Estonia and Ukraine signed their mutual Free Trade Agreement in May 1995,[5], which went into force in January 1997. It provided complete elimination of tariffs and quotas on *all* merchandise trade, including in agricultural products. Additionally, both sides were obligated to not introduce any new tariffs or quotas while the agreement was in force, which created considerable policy certainty. Furthermore, the agreement included important behind–the–border provisions, in particular regarding non–discrimination in public procurement (§9), competition, and intellectual property rights. The implementation of the agreement was overseen through a "Joint Committee" consisting of "equally authorized representatives" of both countries, acting on the consensus principle.

The agreement was terminated by May 1st 2004, when in the course of the EU Eastern Enlargement, Estonia along with seven other Eastern European countries (collectively known as "A8")[6] joined the European Union, which before consisted of 15 countries (the "EU15"). Upon EU accession, Estonian trade policy underwent a discontinuity. Its trade policy changed overnight: while the EU accession granted single-market access to all "A8" countries starting in May 2004, it also demanded that all countries adopt the common EU trade policy, i.e. they enjoyed all benefits of the EU's single market but traded with all non-EU countries at the terms that were negotiated between the EU and those third countries up to May 2004. As a consequence, the Estonian MFN tariff more than doubled (reaching 4.18% by 2005). Moreover, since the EU had no trade agreement with Ukraine in place when the 2004 EU Eastern Enlargement took place, Estonia had to terminate their FTA with Ukraine and apply the EU agreements instead. Therefore, similar to Brexit, the "Uxit" required Estonia and Ukraine to again charge tariffs on imports from the former FTA-trading partner. Figure 1(a) shows the resulting changes to Estonia's preferential trade regime. In particular, note that Ukraine lost

[5] The full text of the agreement is available through the *Global Preferential Trade Agreements* database, see World Bank (1995)

[6] Besides the A8, Cyprus and Malta joined the EU on the same day.

its FTA status because there was no EU–Ukraine FTA in place at the time[7]. Instead, Estonia now applied some preferential tariff reductions for Ukrainian imports based on the EU's *Generalized Scheme of Preferences* (GSP). Russia, on the other hand, was earlier treated as a third country on an MFN basis but now became eligible for GSP preferences in Estonia. Both Estonia and the EU had longstanding FTAs in place with the *European Free Trade Association (EFTA)*, comprising Norway, Iceland, Switzerland, and Liechtenstein. Hence, the status of the EFTA countries did not change after the Estonian EU accession.

In line with the provisions for "denunciation" of the Estonia–Ukraine FTA (§28), Estonia provided notice of termination in October 2003. Hence, by this date already, all market participants could anticipate with certainty that the agreement would end. This might have led to anticipation effects, such as companies shifting their trade relations while they still were benefiting from the preferential FTA treatment. Indeed, visual inspection of the trade flows (see Figure 1 (b)) suggests this to be the case, as the trade volume between Estonia and Ukraine already started decreasing in the second half of 2003.

The Estonia–Ukraine dyad was the only one to suffer a "downgrade" of its trade relations. Overall, the EU Enlargement process shows careful sequencing to avoid disruption of existing trade relations. All accession countries had FTAs with the EU15 and EFTA countries already in place; these trade links either were "upgraded" to the Single Market or stayed in place as before. For Estonia, forgoing the Ukraine FTA was an acceptable loss in economic terms: its imports from Ukraine amounted to €87m per year on average from 1999–2003, accounting for 1.7% of total imports. It was also unavoidable: for the Union to allow an exception to the common trade policy for this FTA would have been legally and administratively challenging,[8] and an EU–Ukraine FTA was not on the political agenda at the time. Because of these factors, one can think of "Uxit" as a *quasi-experiment*, which occurred for reasons entirely unrelated to any bilateral Estonia-Ukraine shocks.

3. Data & Specification

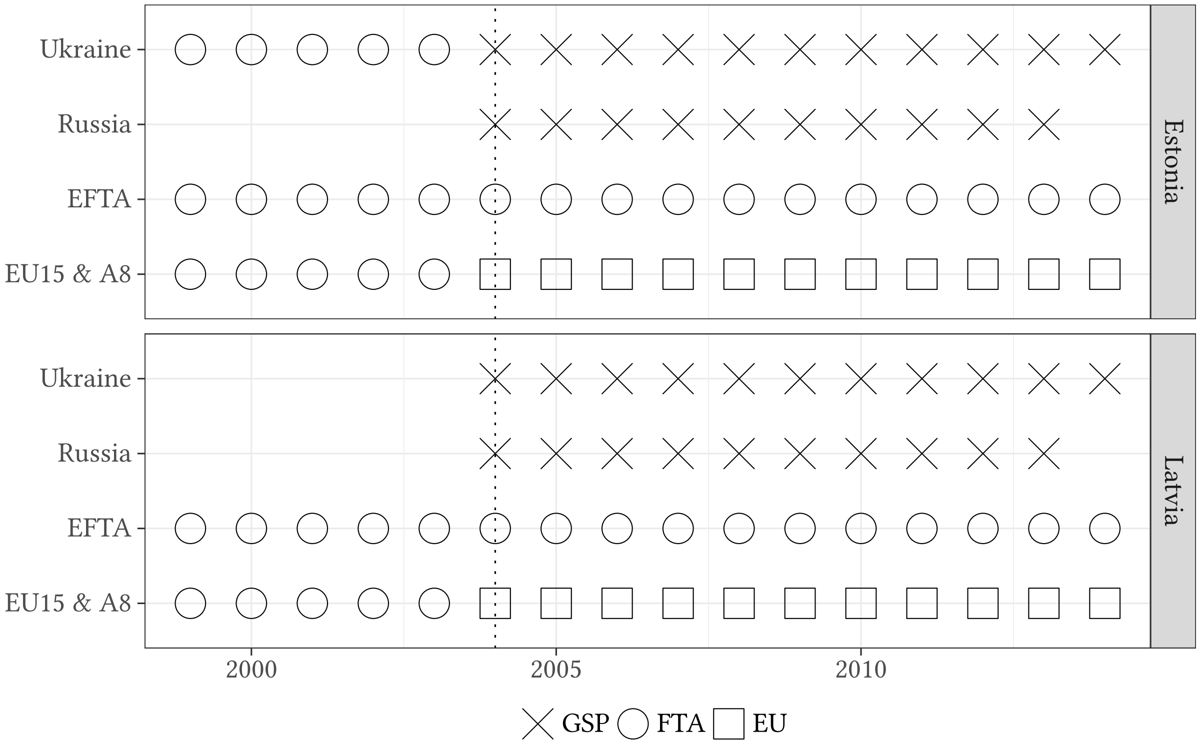
We use this quasi-experiment to estimate the causal effect of the dissolution of a free trade agreement on countries' trade volumes and welfare. For this, we start with triple difference-in-differences OLS panel-regressions, comparing trade-flows between Estonia and Ukraine to two suitable reference countries to see how their trade-flows respond to "Uxit". Afterwards, we proceed with a partial equilibrium analysis and estimate the pooled as well as directional effects of FTA withdrawal using Pseudo-Poisson Maximum Likelihood (PPML) estimation. All

[7]Interestingly, the new "Deep and Comprehensive FTA" between the EU and Ukraine, in effect since 1st January 2015, is less comprehensive in terms of tariff elimination than the earlier Estonia–Ukraine FTA. According to the WTO, various lines are exempted.

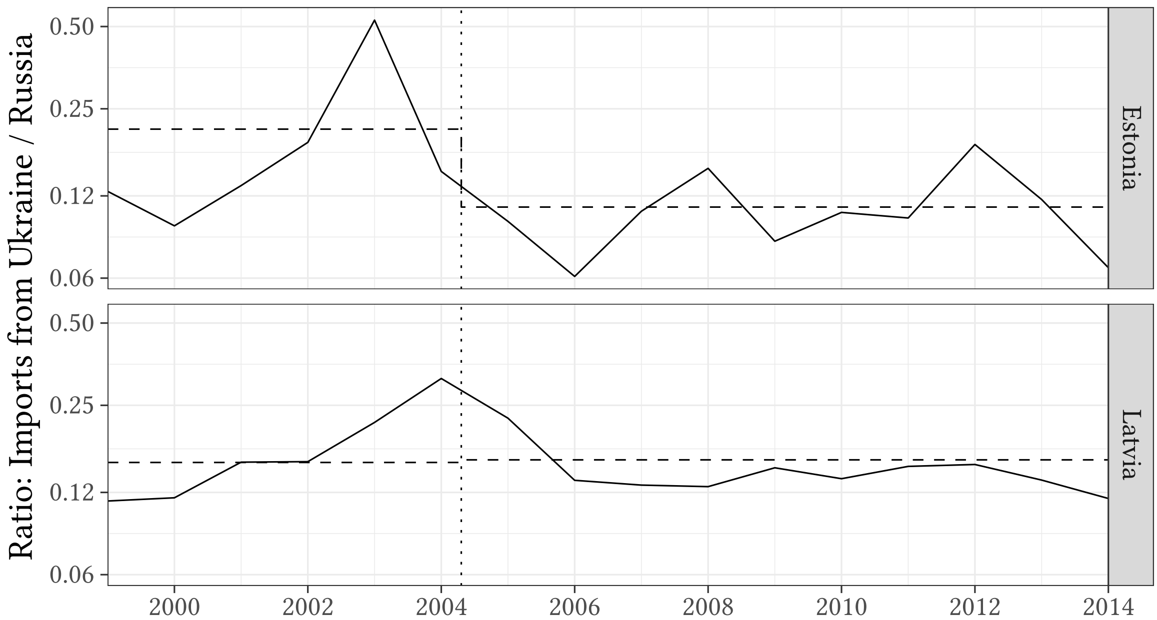
[8]Contrary to theory, there are instances where members of a Customs Union can still have different FTAs. For example, Turkey and the EU are in a Customs Union but have some non-overlapping FTAs (World Bank, 2014)

FIGURE 1
Trade Policy and Trends in the Baltics

(a) Structure of Preferential Trade Regimes



(b) Trends in Import Penetration by Ukraine, Relative to Russia



Notes: The vertical dashed line indicates the timing of the EU Eastern Enlargement in May 2004, the accession of inter alia 8 Eastern European countries ("A8"); the A8 include Estonia and Latvia. EU15 are the 15 member states of the EU prior to the enlargement; EFTA is the European Free Trade Association. Trade agreement data based on the EIA September 2015 database (Baier et al., 2014), updated by the authors after 2012. Besides GSP, the EU and Russia have a "Partnership and Cooperation Agreement" in force since 1997. Import penetration figures are the authors' calculation based on the Eurostat COMEXT database.

estimations are based on the canonical structural gravity model of bilateral trade as surveyed in Head and Mayer (2014).

Model: The functional form of our estimations follows the standard gravity model of international trade. This model, also considered the "workhorse" of the empirical trade literature (Head and Mayer, 2014), decomposes bilateral trade flows X_{ijt} of goods from exporting country j to importing country i in year t into the product of country–specific effects relating to the importer and exporter, as well as effects specific to the individual importer–exporter countries dyad:

$$X_{ijt} = \mathbf{M}_{it}^{IM}\ \mathbf{M}_{jt}^{EX}\ \mathbf{T}_{ij}\ \mathbf{D}_{ijt}\ \mathbf{G}_t\ \eta_{ijt} \tag{1}$$

Here, $\mathbf{M}_{it}^{IM}$ and $\mathbf{M}_{jt}^{EX}$, commonly referred to as countries' "monadic attributes", describe attributes that are specific to any exporter j or importer i, and vary over time t. These attributes hence capture all country-specific characteristics that may affect the trade flows between two countries. One possible monadic effect is a country's GDP, reflecting the idea that bilateral trade depends on the combined "economic mass" of two countries. However, as the seminal contribution of Anderson and Van Wincoop (2003) points out, monadic effects cannot be reduced to observable characteristics like GDP, prices, or resources. Gravity models also require to take into account the "multilateral resistances", i.e. the trade barriers a country faces with *all* other countries.[9] Other than by those monadic attributes, bilateral trade is determined by dyad-specific bilateral attributes, captured by $\mathbf{T}_{ij}$ and $\mathbf{D}_{ijt}$. While $\mathbf{T}_{ij}$ is time-invariant and includes constant characteristics like the geographic distance between two countries or sharing a common colonial history, $\mathbf{D}_{ijt}$ contains dyad-specific characteristics that vary over time, most importantly, tariffs or having a trade agreement in place. Finally, $\mathbf{G}_t$ accounts for a global time-trend in trade that is common to all countries. The error term η_{ijt} captures the remaining dyad-year variation that cannot be explained by observable unilateral or bilateral variables. Our coefficient of interest, the impact of withdrawing the FTA between Estonia and Ukraine on their mutual trade flows, is contained in the time-varying, bilateral characteristics $\mathbf{D}_{ijt}$. We can decompose this bilateral part into its separate components to arrive at the following equation:

$$\mathbf{D}_{ijt} = \beta_1 fta_{ijt} + \beta_2 ftaWithdrawn_{ijt} + \sum_{k \in [2003.1, 2004.12]} \theta_k I(i = EE, j = UA, t = k) \tag{2}$$

[9] Consider the example given by Adam and Cobham (2007): There is a reduction in bilateral trade costs between the UK and France, but no change in France–Italy trade costs. As France now can trade cheaper with the UK, its overall multilateral resistance falls, causing some trade flows being diverted from France–Italy to France–UK.

In Equation 2, the dummy variable fta_{ijt} indicates with a value of one whether at a given time t, the dyad ij trades under terms defined by an FTA. Thus far, the whole gravity specification is completely standard (see e.g. Baier and Bergstrand (2007) for a more detailed discussion). The central interest of our study is to identify the trade effect of FTA *termination.* Therefore, we include another dummy variable, $ftaWithdrawn_{ijt}$, indicating with a value of one whether the dyad ij had been trading under FTA terms at some time $t - s$, but is no more trading under FTA-terms at time t. The coefficient β_2 then measures the effect of "Uxit". In our case, the variable $ftaWithdrawn_{ijt}$ will take the value of one for the Estonia–Ukraine dyad in May 2004 and all months thereafter. In addition, to control for anticipation effects that may affect Estonia–Ukraine trade around the period of the FTA termination, in some specifications, we add a set of dyad-time dummies θ_k for the Estonia–Ukraine dyad for the period immediately before and after the FTA lapsed.

From Equations 1 and 2, we can derive an estimation equation to analyze the effect of the FTA withdrawal between Estonia and Ukraine on their bilateral trade flows. In order to take account of the whole intuition of the gravity model and capture the time-invariant, dyad-specific attributes together with country-year specific attributes including multilateral resistances, we use a triple difference-in-differences estimator (Head and Mayer, 2014).

To arrive at this equation, we start by comparing the trade flows from two distinct exporters, j and k, that arrive in the same importing country i over time. The gravity equation tells us that we can express this relative import penetration by the ratio of the monadic exporter effects (resembling the relative "competitiveness" of the two exporting countries), multiplied with the ratios of dyadic effects and the error terms:

$$R_{i[jk]t} = \frac{X_{ijt}}{X_{ikt}} = \frac{\mathbf{M}_{jt}^{EX}\ \mathbf{T}_{ij}\ \mathbf{D}_{ijt}\ \eta_{ijt}}{\mathbf{M}_{kt}^{EX}\ \mathbf{T}_{ik}\ \mathbf{D}_{ikt}\ \eta_{ikt}} \tag{3}$$

In other words, by applying Equation 1 and putting the exports from j to i in perspective to the exports from k to i, we can eliminate the global time trend $\mathbf{G}_t$ as well as the monadic importer effect $\mathbf{M}_{it}^{IM}$ from the equation, as both attributes are common to both exporting countries at any point in time. That is, according to the gravity model, they should affect imports from j and k equally, so they do not change the relative import penetration in country i. The variation in this ratio over time is driven by relative dyadic effects, which is what we want to capture, and relative competitiveness of the two exporters, which is not observed and needs to be controlled for. To solve this problem, note – as did Romalis (2007) – that the relative import penetration in a *second importing country* l will be driven by the same unobserved relative monadic effects of exporters j and k:

$$R_{l[jk]t} = \frac{X_{ljt}}{X_{lkt}} = \frac{\mathbf{M}_{jt}^{EX}\ \mathbf{T}_{lj}\ \mathbf{D}_{ljt}\ \eta_{ljt}}{\mathbf{M}_{kt}^{EX}\ \mathbf{T}_{lk}\ \mathbf{D}_{lkt}\ \eta_{lkt}} \tag{4}$$

Hence, by looking at a "tetrad" of four countries, we can first derive two difference-in-differences estimates for each importer i and l, where we trace the relative import penetration of the two exporters j and k over time in each of the two importing countries. Then, we can calculate the difference between the two difference-in-differences estimates for each importing country to filter out the unobservable monadic attributes of the two exporters that may change over time, but do so for each of the two importers in the same way. Calculating this ratio and taking logs yields our main estimating equation:

$$r_{[il][jk]t} = ln\left(\frac{R_{i[jk]t}}{R_{l[jk]t}}\right) = \underbrace{\{ln(\mathbf{D}_{ijt}) - ln(\mathbf{D}_{ikt})\}}_{\text{DiD country i}} - \underbrace{\{ln(\mathbf{D}_{ljt}) - ln(\mathbf{D}_{lkt})\}}_{\text{DiD country l}} + v_{[il][jk]t} \tag{5}$$

where the combined error term $v_{[il][jk]t}$ consists of:

$$v_{[il][jk]t} = (\eta_{ijt} - \eta_{ikt}) - (\eta_{ljt} - \eta_{lkt}) \tag{6}$$

Ergo, Equations 5 and 6 tell us that we can estimate the causal effect of the FTA withdrawal between Estonia and Ukraine in a panel dataset by comparing two difference-in-differences estimates, derived for Estonia and another importer, where each compares the imports from Ukraine to those of another exporting country over time. Therefore, to estimate our model, we require a panel data set which covers at least two importers and two exporters; otherwise, Equation 1 suffers from perfect multicollinearity. In economic terms, one needs a *reference importer* similar to Estonia, to control for Ukraine's exporter-time specific effects, and a *reference exporter* similar to Ukraine, to control for Estonia's importer-time specific effects.[10] For the setting of this paper, this approach is especially suitable because good reference countries are available.

Reference Countries: We complete our tetrad of countries by choosing a reference importer vis-à-vis Estonia, and a reference exporter vis-à-vis Ukraine. In both cases, there are obvious choices: Latvia, the southern neighbor of Estonia, shares a rich common history and has simultaneously pursued European integration with Estonia. Regarding trade agreements, the only difference between the two is the Estonia–Ukraine FTA. While Estonia had an FTA with Ukraine, Latvia and Ukraine traded under MFN tariffs (see Figure 1a). Moreover, the business cycles of these two countries move in sync: GDP growth rates are correlated at 94% over the sample period, as Table C1 shows. Thus, Latvia seems to be as good a reference country as one could hope to find. For Ukraine, its eastern neighbor Russia is likewise an obvious choice, as their economies show important similarities. Although relations have been problematic

[10]We focus on exports from Ukraine to Estonia instead of Estonian exports to Ukraine in OLS due to data availability.

lately[11], GDP growth rates between the two were also highly correlated throughout the sample period (90%)[12]. The two countries differ in their trade relations with third countries (for example, Russia has "union state" with Belarus and formed the Eurasian Customs Union from 2012 onwards, i.e. towards the end of the sample period), which could affect the relative multilateral resistance of the two. While the estimation controls for this, it is reassuring that world export growth of the two countries remained highly correlated throughout the sample period (91%). Hence, the baseline estimation involves the Estonia-Latvia-Ukraine-Russia tetrad.

The advantage of this tetradic approach is that high quality monthly data are available, and the risk of mis-specification is minimal; however, this comes at the downside of a possible loss of statistical efficiency and perhaps robustness. To address these concerns, we consider the EFTA[13] countries as an alternative reference exporter for Ukraine. The gravity model tells us that, in principle, it should not matter which reference countries one uses to estimate the triple difference-in-differences Equation 5, subject to some qualifications. As long as the compound error term is not correlated with changes in trade agreements, the estimates should be unbiased. Choosing "similar" reference countries does, however, help efficiency because it reduces the variance of the compound error.[14] Clearly, the EFTA countries are not as closely correlated with Ukraine (e.g. GDP growth correlation is merely 70%), so some loss of efficiency can be expected. Still, it is a reassuring robustness check to find that results do not depend on the details of country choice, but are similar for the Estonia-Latvia-Ukraine-EFTA tetrad. In the Appendix, we provide another robustness check by replacing Latvia by Lithuania as a reference–importer for Estonia. Lithuania, as well, moves very closely to Estonia in terms of GDP and joined the EU together with Estonia and Latvia in 2004 (see Table C2). However, as shown in Figure B3, the relative import penetration of Ukraine vs. Russia in Lithuania moves rather close to zero and does not leave much variation to estimate a significant FTA withdrawal effect in an Estonia-Lithuania-Ukraine-Russia tetrad. Therefore, we prefer Latvia as a reference–importer for Estonia, but still report results for Lithuania in the Appendix.

Estimation by Tetrad: To estimate each of our two tetrads, we plug Equation 2 into Equation 4 for Estonia and Latvia, respectively. Then, taking the ratio of the Estonia-equation over the Latvia-equation following Equation 5, for each tetrad we arrive at our estimating equation:

$$r_{[EE,LV][UA,k]t} = \omega + \beta_{2,EE-UA}\,\text{ftaWithdrawn}_{[EE-UA],t} + \upsilon_{[EE,LV][UA,k]t} \tag{7}$$

[11]In Table 2, we estimate the model excluding the months when the crisis started.

[12]See Table C1

[13]The European Free Trade Association (EFTA) allows its member countries, Iceland, Liechtenstein, Norway & Switzerland, access to the EU Single Market and Schengen Area, though they are no part of the EU Customs Union.

[14]For example, given the high correlation of Ukrainian and Russian export shocks, the difference $\eta_{i[UA]t} - \eta_{i[RU]t}$ may be very small

where $k \in \{\text{RU,EFTA}\}$. The intercept ω contains the relative time–invariant dyad fixed effects $\mathbf{T_{ij}}$ and the trade effect $\beta_{1,EE-UA}$ of the Estonia–Ukraine FTA. Since our sample starts in 1999, two years after the FTA between Estonia and Ukraine was formed, the exact FTA-effect cannot be identified. Ergo, we will compare post-withdrawal to pre-withdrawal trade-flows, while ignoring any differences in trade flows that occurred after signing the FTA.[15] Note further that both importers, Estonia and Latvia, change to GSP tariffs with Ukraine and Russia by entering the EU in May 2004. While Estonia "downgrades" from an FTA to GSP with Ukraine and "upgrades" from MFN tariffs to GSP with Russia, Latvia "upgrades" from MFN to GSP tariffs with both exporters. However, any trade effects of the GSP program are controlled for by our specification, as both importers granted and revoked GSP preferences to both exporters at the same time, so the tetrad difference is nil in each time period. To this baseline model, we add various time dummies for the period around the EU accession to capture possible anticipation effects on trade. The coefficient on the time dummy, which marks the Eastern EU enlargement from May 2004 onwards, then captures the pure effect of "Uxit".

Panel Data: For robustness, we combine the two individual tetrads to a panel data set. With this, we formally test whether the estimates of the FTA withdrawal effect depend on the choice of the reference country. If they do not, then it is possible to use data from both tetrads jointly, while allowing for tetrad–fixed effects, to get a more precise coefficient estimate. To make sure the standard errors are reliable and do not give a false sense of precision, we use clustering similar to Head et al. (2010)[16].

PPML Estimation: Finally, we estimate a structural gravity model with a dyadic panel dataset covering 179 countries and the yearly flows over 1992–2012 using Pseudo-Poisson-Maximum-Likelihood (PPML). In the trade literature, PPML is the best-practice estimator for the gravity model of trade. PPML is advantageous to OLS panel estimates as it allows solving the gravity equation directly in multiplicative form instead of taking logs and hence accurately accounts for the zero-trade flows as well as heteroscedasticity in the trade data (Santos Silva and Tenreyro, 2006). We follow the nomenclature introduced in Baier and Bergstrand (2007) and Baier, Yotov, and Zylkin (2019) and describe trade flows $X_{ij,t}$ between origin i and destination j in year t as:

$$X_{ij,t} = \frac{A_{i,t} w_{i,t}^{-\theta} \tau_{ij,t}^{-\theta}}{P_{j,t}^{-\theta}} E_{j,t} \tag{8}$$

[15] Note: We make the comparison of after-FTA to before-FTA trade flows in the PPML analysis below.

[16] Head et al. (2010) use three-way clustering at the main exporter–year (jt), main importer–year (it) and main importer–main exporter (ij) level. As our estimations contain only one main importer (Estonia) and one main exporter (Ukraine), we use two-way clustering accounting for jt and it.

In this framework, bilateral trade depends on the total expenditures $E_{j,t}$ at destination j as well as quality of production $A_{i,t}$ and wages $w_{i,t}$ at origin i. Additionally, the inward multilateral resistance $P_{j,t}$ accounts for the average import competition at destination j. All these factors vary over time, but are specific to either exporter i or importer j and can therefore be controlled for by including exporter-year and importer-year fixed effects in panel regressions (Anderson and Van Wincoop, 2003). Adding $\mu_{i,t}$ and $\pi_{j,t}$ to account for origin-specific and destination-specific effects as well as an error term $\epsilon_{ij,t}$ to account for unobserved heterogeneity, and re-writing Equation 8 in exponential form, we arrive at:

$$X_{ij,t} = exp\left[\mu_{i,t} + \pi_{j,t} + ln(\tau_{ij,t}^{-\theta})\right] + \epsilon_{ij,t} \tag{9}$$

Here, we are interested in the iceberg trade costs $\tau_{ij,t}$, which we can linearize as

$$ln(\tau_{ij,t}) = Z_{ij}\delta + \beta_1 FTA_{ij,t} + \beta_2 FTAWithdrawn_{EE-UA,t} + u_{ij,t} \tag{10}$$

Following the insights from Baier and Bergstrand (2007) and as recommended by Baier, Yotov, and Zylkin (2019), we add dyad-fixed effects χ_{ij} to control for dyad-specific effects Z_{ij} that do not vary over time. We further include an indicator variable for an FTA being in place between countries i and j in year t, therefore estimating the coefficient β_1. Finally, we add the dummy variable $FTAWithdrawn_{EE-UA,t}$ to our regressions, which takes the value of one for the Estonia–Ukraine dyad for all years from 2004 onwards. Subject to the control variables and fixed effects included in our regressions, and based on our discussion above about "Uxit" being exogenous to the bilateral relationship between Estonia and Ukraine at the time of Estonia's EU accession, we can assume that our Withdrawal dummy and the error term $u_{ij,t}$ are uncorrelated. Hence, we can interpret β_2 as the causal effect of FTA Withdrawal on bilateral trade between Estonia and Ukraine. Concluding, our estimation equation can be written as follows:

$$X_{ij,t} = exp\left[\mu_{i,t} + \pi_{j,t} + \chi_{ij} + \beta_1 FTA_{ij,t} + \beta_2 FTAWithdrawn_{EE-UA,t}\right] + u_{ij,t} \tag{11}$$

where we cluster the composite error term $u_{ij,t}$ at the dyad-level. As further variants, we include different dummy-variables which indicate a one-way change in the trade relationship instead of the dyadic view, e.g. that Estonia has an FTA with Ukraine in a given year. Compared to the tetradic estimations discussed above, these PPML regressions provide another way to account for observable and unobservable monadic effects, drawing on information from more than 700,000 observations. Further, it allows us to look into directional effects and see whether the effect of the FTA withdrawal was larger for Estonia or Ukraine. Signing an

FTA does not necessarily affect all signatories similarly (Baier, Yotov, and Zylkin, 2019), which lets us expect that the same might be true for the withdrawal of an FTA.

Data: For the OLS estimations, we use data on bilateral trade flows from the Eurostat "Comext" database, which covers import flows reported by the Baltic countries of Estonia and Latvia, as well as the aggregate of the "EU15" countries. For each reporter, we collect total imports by all partner countries. The data are reported monthly, which allows us to trace the immediate trade-response to the EU enlargement and associated trade policy "treatments", which took place on May 1st 2004, i.e. in the middle of the year. The sample period runs from January 1999, the first year for which data are available for the Baltic countries, to December 2014.

For our PPML estimations, we use annual bilateral world trade data based on the COMTRADE database covering 179 countries for the period 1992–2012. We include a measure for bilateral trade relations as a control variable, which is based on the September 2015 version of the *Economic Integration Agreements* database of Baier et al. (2014), updated through to 2014 by the authors.

4. Results

We begin by providing OLS results as a benchmark, first from tetrad regressions and second from the combination of the two tetrads to a monthly panel dataset. Then, we follow the recommendation of Santos Silva and Tenreyro (2006) and Santos Silva and Tenreyro (2011) to estimate the model using Poisson Pseudo-Maximum Likelihood (PPML) using annual global trade data. Finally, we provide estimates of welfare effects by (a) using our PPML estimates for the withdrawal effect to compute counterfactual trade flows and welfare levels for all sample-countries had "Uxit" not taken place, and (b) comparing these counterfactual trade flows and welfare levels to the actual, observed ones.

Tetrad Regressions: Table 1 presents the results from our first estimation method using the tetrad-estimation OLS approach[17] described by Equation 7. Columns (1)–(3) report our results from using Russia as a reference exporter to Ukraine, while Columns (4)–(6) report the results for the four EFTA–countries as reference exporter. Throughout all columns, the coefficient estimate for our post-May 2004 Dummy variable indicating the resolution of the FTA between Estonia and Ukraine is negative and statistically highly significant, suggesting that FTA withdrawal leads to a noticeable fall in trade.
Model (1) reproduces the trade pattern already shown in Figure 1(b) in Section 2, by estimating the gross difference in Estonia's relative import penetration by Ukraine relative to Russia, vis-á-vis the relative import penetration in Latvia. Model (2) adds dummies for the years

[17] We reproduced the same regressions using PPML instead of OLS. The results look very similar (not reported).

TABLE 1
Tetradic Regression Results

	(1)	(2)	(3)	(4)	(5)	(6)
ftaWithdrawn	-0.463^{***}	-0.279^{***}	-0.261^{***}	-0.502^{***}	-0.273^{**}	-0.256^{**}
	(0.113)	(0.089)	(0.095)	(0.137)	(0.116)	(0.124)
Reference Exporter	RU	RU	RU	EFTA	EFTA	EFTA
Controls						
Years '03, '04		X	X		X	X
Months in '03, '04			X			X
Durbin-Watson	0.00	0.00	0.00	0.00	0.00	0.00
Breusch-Pagan	0.71	0.06	0.82	0.13	0.00	0.45
Jarque-Bera	0.13	0.66	0.79	0.32	0.58	0.66
Observations	192	192	192	192	192	192
R^2	0.124	0.301	0.360	0.093	0.221	0.253

Notes: Newey-West HAC standard errors in parentheses. *p<0.1; **p<0.05; ***p<0.01. Results from OLS regressions with the logarithm of the ratio of relative import penetrations as dependent variable. The unit of observation is the Tetrad-month from January 1999 to December 2014. The main explanatory variable is a dummy variable indicating months after May 2004. In the year controls specification, indicators are added for the year being 2003 or 2004. In the month controls specifications, additional indicators are added for each month from February to December in the years 2003 and 2004. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text.

2003 and 2004, which let us control for anticipation effects that may have occurred during the period between the announcement of the FTA withdrawal and the de facto withdrawal.[18] This results in a point estimate of -0.279, which suggests a trade reduction of around 24.3% due to the FTA withdrawal.[19] In Model (3), we add further dummies for each month from January 2003 to December 2004, which effectively causes the observations from this period to drop out from our OLS estimation so that we basically compare the ratio-of-ratios in trade flows from January 2005 and later to the period of 1999 to December 2002. The inclusion of these monthly dummies hardly has any effect on the coefficient estimate, which decreases slightly to -0.261 but retains its economic and statistical significance. Although this specification has a minimally higher R^2 than Model (2), a Wald test gives no evidence of an improved fit of the model ($F_{166,22} = 0.7047$). This lets us conclude that annual dummies are enough to capture the anticipation effect for the Russia tetrad. The effect we estimate through Columns (1)–(3) is economically important and statistically significant at the 1% level throughout. Our most conservative estimate from Model (3) points towards a trade-reduction of around 23%,

[18]From Figure 1(b), one might expect an anticipation effect already setting in at the beginning of 2002. In Table C4, we replicate Table 1 including dummies for the year 2002. Despite a great loss in power from essentially dropping another 12 months from our sample, the results are robust for Russia as a reference exporter, but slightly fall short of statistical significance for the EFTA countries.

[19]Following the equation $(e^{-0.279} - 1) \times 100\% = 24.3\%$.

suggesting that an FTA withdrawal mostly eliminates all gains from signing an FTA in the first place, which has been estimated to lie around 20.7% on average (Baier, Yotov, and Zylkin, 2019). Further, our estimates exceed the simulated effects of a "Hard Brexit" on UK-EU trade, which lie around 12% (Felbermayr, Gröschl, and Steininger, 2018).

In Models (4–6), we repeat the estimations outlined by Equation 7, but substitute the EFTA countries as the reference exporters for Russia. Reassuringly, the point estimates are very close in size to models (1–3) and remain statistically significant at least at the 5-percent level. As expected, the estimates are slightly less precise when EFTA is used as the reference exporter (the standard errors increase by about 1/3), suggesting that the monadic shocks of Ukraine are more correlated with the monadic shocks of Russia than with those of the EFTA countries. Again, monthly dummies for the transition period do not improve the fit of the model compared to the annual dummies, but reduce the precision of the estimates. Therefore, for the estimations to follow, we will stick to yearly dummies according to Models (2) and (5) and no more report estimations using monthly dummies.

The residuals appear to be normally distributed according to the Jarque-Bera statistic throughout. The Breusch-Pagan test scores however report clear evidence for heteroscedasticity in the EFTA model with year-dummies (5), which appears to be driven by more volatile EFTA trade shares in both Estonia and Latvia towards the end of the sample period. There is also weak evidence for heteroscedasticity in Model (2), using Russia as the reference exporter and also adding year-dummies. Further, the Durbin-Watson statistic provides strong evidence for serial correlation across all our specifications. Visual inspection of the residuals for Model (2) clearly shows "runs" of positive and negative residuals across adjacent months (see Figure B7). There were several months with positive residuals, e.g. in 2002 and 2008, and runs with negative residuals in e.g. 2001 and 2005. The picture looks similar for the EFTA tetrad (Model 5), though autocorrelation seems to be even more of a problem there. Between 2006 and 2008, the residuals constantly remained in the positive segment.[20]

Since we are using monthly data, one possible source of autocorrelation is seasonality. However, the correlogram in Figure B7(b) shows that autocorrelation is significant only up to 6 months. At month 12, the hypothesis of no autocorrelation cannot be rejected at the 5% level, suggesting that the tetrad design successfully solves any seasonality issues. For the EFTA countries however, the autocorrelation appears significant even after 15 months, so seasonality cannot be ruled out. Therefore, we re-run our estimations of Table 1 including month-of-the-year fixed effects, which leaves our results unchanged (see Appendix Table C5). These findings suggest that serial correlation is driven by short-run adjustment dynamics rather than misspecification of the model.

Nevertheless, given the undeniable presence of autocorrelation, OLS may not be an efficient estimator. Thus, we re-estimate our preferred Models (2) and (5) using the Cochrane-

[20] See Figure B8 in the Appendix.

Orcutt estimator, which adjusts the model for AR(1) errors. We report the results in the first two columns of Table 2, denoted Models (7)–(8). Using the Cochrane-Orcutt estimator instead of OLS increases both the point-estimate as well as the error terms a bit. While for our Russia–tetrad, the negative effect remains significant at the five percent level, the estimate slightly misses the threshold for 10-percent significance for the EFTA–tetrad. Apparently, adjusting for autocorrelation in the standard error does not leave enough monthly variation in the EFTA-sample to identify a significant effect. Therefore, we turn to our panel estimations where we pool the two tetrads in our sample, hence doubling the number of observations and therefore statistical power, while further controlling for autocorrelation in the error terms.

TABLE 2
Further Regression Results

Estimator	Cochrane-Orcutt		Panel, Fixed Effects			
	(7)	(8)	(9)	(10)	(11)	(12)
ftaWithdrawn	-0.285^{**}	-0.274	-0.276^{**}	-0.257^{**}	-0.239^{**}	-0.350^{**}
	(0.133)	(0.186)	(0.124)	(0.102)	(0.106)	(0.169)
ftaWithdrawn×EFTA				-0.039	0.040	0.040
				(0.096)	(0.096)	(0.097)
ftaWithdrawn×Year<2009						0.228
						(0.255)
Reference Exporter	RU	EFTA	Both	Both	Both	Both
Controls: Years ’03, ’04	X	X	X	X	X	X
Observations	192	192	384	384	358	358
Adj. R^2			0.267	0.266	0.282	0.296

Notes: In panel estimation, clustered Newey-West HAC SE are used to control for heteroscedasticity and autocorrelation in the error term. *p<0.1; **p<0.05; ***p<0.01. Results from OLS regressions with the logarithm of the ratio of relative import penetrations as dependent variable. The unit of observation is the Tetrad-month from January 1999 to December 2014. The main explanatory variable is a dummy variable indicating months after May 2004. In the year controls specification, indicators are added for the year being 2003 or 2004. In the month controls specifications, additional indicators are added for each month from February to December in the years 2003 and 2004. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text. For models (11-12), the sample period is limited from January 1999 to November 2014 only (considered the onset of the “Ukrainian crisis”).

Panel Data Regressions: In Models (9)–(12) of Table 2, we combine both tetrads to a pooled panel dataset and run the same ratio-of-ratios estimation as in Models (1)–(8) before. All models include a tetrad-fixed effect to only trace differences in each tetrad’s ratio, instead of comparing between-tetrad changes. We further apply two-way clustering following Head et al. (2010), clustering standard errors at the it and jt level. Further, we report autocorrelation- and heteroscedasticity-robust standard errors throughout.

Model (9) displays the baseline results, repeating our main regressions with the panel dataset. Reassuringly, the point estimate for the FTA withdrawal lies in between the point-estimates from either tetrad-sample and is statistically significant at the five percent level. From Model (10) onwards, we include an interaction term between the FTA withdrawal dummy and our tetrad-fixed effect to check whether the withdrawal effect is different between the two tetrads. As expected, the interaction term is far from statistical significance throughout all models, confirming our prior that the choice of the reference exporter does not change the qualitative results. Model (10), using the full sample and adding the interaction with the tetrad-fixed effect constitutes our preferred specification. The estimate here suggests that "Uxit" reduced trade between Estonia and Ukraine by about 22.7% on average, and that this effect is not different across tetrads.

In Models (11) and (12), we commence with two robustness tests to our preferred specification. First, in Model (11) we reduce our dataset to the subsample of months before November 2013, which is considered the starting point of the Ukrainian crisis. In theory, the Ukrainian crisis is a monadic shock that our identification strategy should be controlling for. This also seems to be the case in practice, as the estimate is hardly sensitive to the changed sample period. Though the reduction in observations hurts the precision of our estimate a little, it remains statistically significant at the five percent level, while the point estimate only decreases slightly. Lastly, Model (12) adds an interaction, where we additionally interact our FTA withdrawal dummy with a dummy indicating the period before 2009. The idea is to split the post–FTA period into two halves and test the "structural stability" of the coefficient estimates. There is no evidence that the effect was different in the short or medium run. The point estimate decreases to -0.35, but is not statistically different from the point estimate in Model (11). Further, the newly introduced interaction term is far from statistical significance, so we see no evidence that the effect may be different between the early and the later period.

PPML Regressions: Our OLS results provide robust evidence for a negative effect of FTA withdrawal on bilateral trade, relying on three different reference countries. Now, we turn to PPML estimations using yearly data for most countries of the world, and specifying our gravity model in levels rather than ratios. As this is closer to the general method of estimating the effects of signing FTAs, our PPML results provide a better comparison to the trade-enhancing effects of FTAs found in the literature.

Table 3 provides the results of our PPML estimates, where we use bilateral trade data for 179 countries and the years 1992–2012. Following the conventions of the trade-literature, we regress the bilateral trade volume on a bilateral explanatory variable, i.e. the formation or withdrawal of the FTA between Estonia and Ukraine, while controlling for importer-year, exporter-year, and importer-exporter fixed effects and clustering standard errors at the dyad-level. This means, the variation we are analyzing comes only from shocks that are specific to some country-dyad and which vary over time. In Column (1) of Table 3, we connect to the results from our OLS regressions as a benchmark and restrict the sample-period to years after

TABLE 3
PPML Comtrade Results

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sample:	After 1997	All	All	All	All	All	All
FTA Withdrawn							
Both	-0.672***	-0.551***				-0.230	
	(0.149)	(0.134)				(0.209)	
Exporter = UA			-0.611***				-0.231
			(0.149)				(0.271)
Exporter = EE			-0.469**				-0.245
			(0.210)				(0.303)
FTA in tact							
Both				0.583***		0.362**	
				(0.127)		(0.156)	
Exporter = UA					0.656***		0.436**
					(0.129)		(0.197)
Exporter = EE					0.483**		0.245*
					(0.194)		(0.134)
Observations	303147	369735	369735	369735	369735	369735	369735

Notes: Standard errors clustered at dyad–level in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Results from PPML Regressions with trade volumes as dependent variable. The unit of observation is the dyad-year level for 179 countries and the years 1992 to 2012. The main explanatory variables are dummies that indicate either the conclusion or withdrawal of the FTA between Estonia and Ukraine. We alternate between non-directional (Columns 1, 2, 4, 6) and directional dummy variables (Columns 3, 5, 7). All regressions include exporter-year, importer-year as well as dyad Fixed Effects and control for the existence of trade agreements between countries, i.e. FTAs, PTAs, Currency Unions, Common Markets or common EU-membership.

1997, when the FTA between Estonia and Ukraine went into force. Hence, the regression in Column (1) compares trade between Estonia and Ukraine during the after-FTA period to the period during the FTA was active. We find a negative effect of "Uxit" which is both statistically and economically highly significant. On average, trade between the two countries was around 47% lower after the FTA was withdrawn, compared to the FTA-period. Note that our regressions control for all trade agreements other than the Estonia–Ukraine one following Baier and Bergstrand (2007). Column (2) extends the sample and also includes the years before the FTA was signed. Hence, we now compare the post-FTA period to the years when the FTA was in tact as well as the years before the two countries signed their FTA. For this full sample, the point estimate decreases in size as was to be expected, but retains its significance. Both columns hence confirm our earlier findings from the OLS regressions. In Column (3), we take a look at the directional effects of the FTA withdrawal following Baier, Yotov, and Zylkin

(2019). This is, instead of regressing trade volumes on one dummy variable that indicates the Estonia-Ukraine dyad in years later than 2004, we include two separate dummies. One dummy indicates export flows from Ukraine to Estonia after 2004, the other dummy variable takes the value of 1 for exports from Estonia to Ukraine after 2004. We see here that the exporters from both countries significantly reduced their exports to the partner country after the FTA was resolved. Even though the point estimate is larger for the Ukraine-Estonia connection than for the Estonia-Ukraine connection, an F-Test rejects a significant difference between the two estimates. Hence, the FTA withdrawal affected both countries' exporters similarly, ruling out differences in directional effects.

Starting in Column (4), we introduce dummy variables for the FTA being active. This exercise is meant to test whether the Estonia–Ukraine FTA constitutes a special case, or matches the previous findings in the trade literature. Again, we distinguish between pooled effects in Column (4) and directional effects in Column (5). We find that on average, trade between the two countries was around 82% higher during the FTA was in tact. According to column (5), even though the point estimate for the Ukraine–Estonia direction looks bigger, an F-Test again rejects a significant difference between the two estimates and suggests that the effect was not directional. These results slightly exceed the range of earlier studies, which find an average positive FTA-effect on trade between 26% and 58% (Baier, Yotov, and Zylkin, 2019; Baier and Bergstrand, 2007). This difference probably is due to the fact that Ukraine constituted an internationally rather closed country, especially during the late 1990s and early 2000s. Hence, the free-trade regime with Estonia likely had an exceptionally large effect on Ukrainian consumers as well as exporters, who faced high tariffs for imports from and exports to most other countries.

Finally, we add both the FTA-withdrawal and the FTA-in-place dummies together, first un-directional in Column (6), then as directional dummies in Column (7). By including both types of dummies at the same time, we hence compare the FTA-period and the FTA-withdrawal-period to the time before the FTA between Estonia and Ukraine was signed. This exercise confirms the findings for the FTA-accession effect. Trade was on average 43% higher during the FTA than in the years before.

Further, an F-Test again rejects a significant difference between the directional effects. Additionally and most interestingly, the point-estimates for the withdrawal-dummies, both directional and un-directional, turn insignificant. This suggests that the withdrawal of the FTA between Estonia and Ukraine has almost exactly undone the gains from signing the FTA. According to our results, there is no significant difference in trade values between Estonia and Ukraine in the period before and the period after the FTA was in place. We interpret this finding as evidence that no "institutional memory" is left behind when two countries cancel their FTA.

General Equilibrium Effects: As a final exercise, we investigate how "Uxit" affected the

TABLE 4
Welfare Changes from FTA Withdrawal and EU Accession, in Percent

Country	Welfare Change
Latvia	0.021
Lithuania	0.015
Finland	0.007
Russia	0.003
Bulgaria	0.002
Cyprus	0.002
Egypt	0.002
Sweden	0.002
Ukraine	-0.135
Estonia	-1.175

(a) FTA Withdrawal Only

Country	Welfare Change
Slovenia	0.938
Lithuania	0.894
Czech Republic	0.872
Hungary	0.869
Latvia	0.856
Cyprus	0.547
Poland	0.527
Estonia	0.224
Austria	0.096
Ukraine	-0.159

(b) Including EU Accession

Notes: General Equilibrium calculations based on PPML estimates from a symmetric dataset covering 56 countries from 1999 – 2016. Numbers show the computed welfare difference (in percent) of the actual world compared to counterfactual scenarios. Panel (a) computes these changes for a counterfactual world where the FTA between Estonia and Ukraine was never withdrawn. Panel (b) also accounts for the FTA withdrawal, but additionally incorporates the GE effects of the EU accession round in 2004.

welfare levels of Estonia and Ukraine, as well as other third countries. To do so, we create a yearly dyadic dataset similar to the one we used for our PPML regressions, but do a couple of adjustments following Baier, Yotov, and Zylkin (2019). Most importantly, we extend our dataset by internal trade flows and reduce our dataset to only include symmetric trade flows from all exporters to all importers in every year.[21] This results in a dataset that contains 41 countries and covers the years 1998–2009.

With this dataset, we first re-run our main partial equilibrium PPML regression following Equation 11. The coefficient estimates are similar to the ones found in Table 3, although the point estimates are a little larger in their absolute value (see Appendix, Table C8). While the estimates are not significantly different from the ones in Table 3, the bigger size likely stems from the focus on manufactured goods[22].

We use in the next step the estimate from Table C8, Column (1), to compute general equilibrium changes in trade following a one sector Armington-CES model, assuming a constant trade elasticity of $\theta = 4$.[23] In essence, this gives us counterfactual trade flows as well as coun-

[21] Appendix D provides more details on the dataset construction.

[22] Theoretically, also the inclusion of internal trade flows can lead to bigger point estimates. However, running the same regressions with the same sample, but excluding internal trade flows, yield very similar results (not shown).

[23] We use the "ge_gravity" Stata Command provided by Thomas Zylkin based on Baier, Yotov, and Zylkin (2019) for this task.

terfactual welfare levels for a world in which Estonia would not have withdrawn from its FTA with Ukraine. Welfare is defined as total national expenditure relative to the price level. Now, we can compare this counterfactual welfare and trade levels to the actual, observed ones.

Table 4, Panel (a) shows the results for the main affected countries. Unsurprisingly, Estonia and Ukraine faced the largest decreases in welfare from the FTA withdrawal. Whereas the effect on Ukrainian welfare is already large with a decrease of around 0.14%, the effect is even more remarkable for Estonia. Had the FTA with Ukraine not been withdrawn, Estonia's level of welfare would have been almost 1.2% higher in the years afterwards. Those two main countries aside, there are several third countries that benefited from the FTA withdrawal, likely due to trade diversion effects. Especially Latvia and Lithuania benefited with welfare increases of around 0.02%.

Panel (a) only paints half of the picture for Estonia, though. Whereas the FTA withdrawal with Ukraine certainly hurt domestic exporters, both exporters and consumers at the same time enjoyed the accession to the EU. Therefore, in Panel (b) we simultaneously account for both events that occurred to the European economies in 2004: the withdrawal of the FTA between Estonia and Ukraine, *as well as* the A8 countries joining the European Union. Arguably, this provides a better picture of the effects of EU accession.

We find first and foremost that the A8 countries benefited largely from the EU accession[24]. Slovenia and Lithuania lead the list with a welfare increase of around 0.9%. Additionally, we find that Austria also enjoyed a 0.1% welfare increase, even though it was already part of the EU in 2004. This likely stems from trade benefits with its neighboring countries Slovenia, Slovakia, Hungary and the Czech Republic which all joined the EU.

Most importantly, Estonia overall enjoyed a welfare *increase* from its EU accession; this is, the benefit of facilitated trade with the other EU countries more than offset the decreased welfare from the FTA withdrawal with Ukraine. Still, its welfare increase of 0.224% is noticeably lower than the benefits enjoyed by the other A8 countries who did not have to sacrifice any trade agreement in order to join the EU.

Finally, we find that once both effects are accounted for, the negative welfare effect for Ukraine becomes even more severe, increasing in size to a welfare loss of almost 0.16%. This finding suggests that Ukrainian consumers not only had to face a loss in welfare due to the FTA withdrawal with Estonia, but also due to trade diversion effects as the newly accessed A8 countries turn more towards trade with other EU countries.

To trace the roots of the estimated welfare effects, we plot the most significant estimated trade changes in Figure B9 in the Appendix. To little surprise, bilateral trade between Estonia and Ukraine decreased the most; exports in both directions decreased by around 64% due to the FTA withdrawal. Trade diversion effects however led to increased exports to other

[24]The only A8 country missing on the list is Slovakia, which we had to drop from our sample due to estimated negative internal trade flows.

destinations. Both Estonia and Ukraine increased their internal trade flows (by 3% and 0.3%, respectively) as well as their shipments to Russia (by 2% and 0.116%, respectively). Estonia further sent more exports to its direct neighbors and the Scandinavian countries, especially to Sweden and Finland.

5. Extensions and Robustness

The main results presented in the prior section outlined a sizeable negative effect on the trade between Estonia and Ukraine due to Estonia's withdrawal from their bilateral Free Trade Agreement in 2004. In this section, we provide some further robustness checks.

TABLE 5
Robustness: Russia Exports

	(1)	(2)	(3)	(4)	(5)	(6)
ftaWithdrawn	0.463***	0.279***	0.261***	−0.039	0.006	0.005
	(0.089)	(0.086)	(0.091)	(0.104)	(0.111)	(0.121)
Reference Exporter	UA	UA	UA	EFTA	EFTA	EFTA
Controls						
Years ’03, ’04		X	X		X	X
Months in ’03, ’04			X			X
Durbin-Watson	0.00	0.00	0.00	0.00	0.00	0.00
Breusch-Pagan	0.71	0.06	0.82	0	0	0.21
Jarque-Bera	0.13	0.66	0.79	0.14	0.02	0.01
Observations	192	192	192	192	192	192
R^2	0.124	0.301	0.360	0.001	0.029	0.051

Results from OLS regressions with the logarithm of the ratio of relative import penetrations as dependent variable. The unit of observation is the Tetrad-month from January 1999 to December 2014. The main explanatory variable is a dummy variable indicating months after May 2004. In the year controls specification, indicators are added for the year being 2002, 2003 or 2004. In the month controls specifications, additional indicators are added for each month from February to December in the years 2002, 2003 and 2004. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text.

Table 5 displays the results for the same tetradic regressions as used in Table 1, but with Ukraine substituted by Russia as the exporter. This means, Table 5 estimates how the trade volumes between Estonia and Russia differed before and after the FTA between Estonia and Ukraine was withdrawn. Columns (1)–(3) present the results using Ukraine as the reference-exporter. Unsurprisingly, the estimates are exactly the same as in Table 1, but with positive instead of negative signs (and with slightly lower standard errors). Ergo, the trade-flows between Estonia and Russia increased vis-á-vis the trade-flows between Estonia and Ukraine as much as Estonia's trade-flows with Ukraine decreased vis-á-vis Russia. More interesting are

Columns (4)–(6). There, we use trade with the four EFTA-countries as a reference for trade with Russia, in effect cancelling Ukraine completely out of our tetrad. Neither of the estimates reported in Columns (4)–(6) is significant and they even switch signs across specifications. However, the results derived from the EFTA-sample again suffer under heteroskedasticity if the month-dummies are not included. Furthermore, the Jarque-Bera test statistic raises concern of a correct distribution of the error terms. Overall, we take this as evidence that Estonia's withdrawal from the FTA with Ukraine negatively affected Estonia's trade with Ukraine, but did not have any effect on Estonia's trade with Russia, our preferred reference-exporter.

TABLE 6
Robustness: Alternated Timing

	(1)	(2)	(3)	(4)	(5)	(6)
ftaWithdrawn	-0.133	-0.146	-0.261^{***}	-0.261^{***}	-0.152^{*}	-0.010
	(0.112)	(0.098)	(0.091)	(0.091)	(0.090)	(0.087)
Withdrawal Year	2001	2002	2003	2004	2005	2006
Reference Exporter	RU	RU	RU	RU	RU	RU
Durbin-Watson	0.00	0.00	0.00	0.00	0.00	0.00
Breusch-Pagan	0.93	0.79	0.82	0.82	0.97	0.96
Jarque-Bera	0.98	0.93	0.79	0.79	0.94	1
Rainbow	0.37	0.38	0.48	0.48	0.14	0.07
Observations	192	192	192	192	192	192

Notes: Newey-West HAC standard errors in parentheses. *p<0.1; **p<0.05; ***p<0.01. All estimations include indicators for the years 2003 or 2004 as well as indicators for each month from February to December in the years 2003 and 2004 as controls. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text.

In Table 6, we alternate how we code the timing of the withdrawal of the FTA between Estonia and Ukraine, and again repeat our main regressions presented in Table 1. Column (4) presents the original estimate, using the actual withdrawal-date of May 1st, 2004. Walking from Column (4) to the left, we predate the withdrawal by one year each, while we postpone it by one year for each Column walking to the right from Column (4). Coding the withdrawal in year 2003 (the year the withdrawal was announced, but not yet enacted) gives exactly the same result as coding the correct year. This similarity occurs by construction, since we "cancel out" the years 2003 and 2004 of our sample via year-dummies to account for a possible anticipation effect. However, coding the date of withdrawal even earlier than 2003 or alternatively later than 2004 (Columns (1)–(2) and (5)–(6), respectively) lets the point estimates become insignificant, as one would expect. This suggests that the trade decrease occurred at the exact timing of the FTA withdrawal and cannot be attributed to a general downward trend around the accession year.

TABLE 7
Extension: Excluding Transport Sector

Estimator	Tetrad				Panel, Fixed Effects		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ftaWithdrawn	-0.139^{**}	-0.150^{**}	-0.065	-0.063	-0.144	-0.149	-0.255^{*}
	(0.070)	(0.074)	(0.096)	(0.103)	(0.092)	(0.100)	(0.132)
ftaWithdrawn×EFTA					0.075	0.150	0.150
					(0.131)	(0.137)	(0.128)
ftaWithdrawn×Year<2009							0.217
							(0.282)
Reference Exporter	RU	RU	EFTA	EFTA	Both	Both	Both
Controls: Years ’03, ’04	X	X	X	X	X	X	X
Estimation	Tetrad	Tetrad	Tetrad	Tetrad	Panel	Panel	Panel
Observations	192	192	192	192	384	358	358
R^2	0.020	0.074	0.002	0.006	0.096	0.121	0.143

Notes: Newey-West HAC standard errors in parentheses. *p<0.1; **p<0.05; ***p<0.01. Results from OLS regressions with the logarithm of the ratio of relative import penetrations as dependent variable. The unit of observation is the Tetrad-month from January 1999 to December 2014. The main explanatory variable is a dummy variable indicating months after May 2004. Indicators are added for the year being 2003 or 2004. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text.

As Figure B2 shows, the peak in exports from Ukraine to Estonia in 2003 was mainly driven by the product-group“Transport Equipment And Parts And Accessories Thereof” as defined under product-code BEC 521. Even over most of the sample period, Estonian imports from Ukraine were mainly dominated by this one sector. Therefore, as an extension we repeat our main regressions leaving out this main sector from our sample and focusing on the other five sectors instead.

The results are presented in Table 7, where we re-produce our preferred models from Tables 1 and 2, excluding imports of the transport sector. The results show that the point estimate for FTA withdrawal becomes significantly smaller. It drops to -0.15 for the Russia-tetrad, but retains its 5-percent significance. For the tetrad regressions using the four EFTA countries as a reference exporter, the effect becomes imprecisely measured and is no more significantly different from zero. Also, the R^2-statistic drops significantly, indicating that without this main sector, our model lacks explanatory power. A similar problem arises in the panel estimations, where the point-estimate for FTA withdrawal drops in statistical significance and just misses the threshold for the five-percent level of significance, though the magnitude of the effect comes close to the estimate for the all–sectors sample when we restrict the panel to the pre-November 2013 period and add the interaction for observations before 2009. We conclude from this exercise that on the Ukrainian side, the shock concentrated on the main-export sector, while the effect for the other sectors was, at best, rather short-lived.

TABLE 8
Extensions PPML

Sample	World		EU 27 only			
	(1)	(2)	(3)	(4)	(5)	(6)
$Withdrawn$	-0.427**		-0.677***		-0.516***	
	(0.169)		(0.110)		(0.119)	
$Withdrawn \times Year > 2008$	-0.263*				-0.333***	
	(0.154)				(0.0747)	
$Withdrawn_{Exporter=UA}$		-0.571***		-0.673***		-0.532***
		(0.141)		(0.149)		(0.162)
$Withdrawn_{Exporter=EE}$		-0.229		-0.682***		-0.498***
		(0.215)		(0.105)		(0.113)
$Withdrawn_{Exporter=UA} \times Year > 2008$		-0.0839				-0.286***
		(0.0724)				(0.0809)
$Withdrawn_{Exporter=EE} \times Year > 2008$		-0.503***				-0.396***
		(0.0871)				(0.104)
Observations	369735	369735	15694	15694	15694	15694

Notes: Standard errors clustered at dyad–level in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Results from PPML Regressions with trade volumes as dependent variable. The unit of observation is the dyad-year level for 179 countries and the years 1992 to 2012. The main explanatory variables are dummies that indicate either the conclusion or withdrawal of the FTA between Estonia and Ukraine. We alternate between non-directional (Columns 1, 3, 5) and directional dummy variables (Columns 2, 4, 6). All regressions include exporter-year, importer-year as well as dyad Fixed Effects and control for the existence of trade agreements between countries, i.e. FTAs, PTAs, Currency Unions, Common Markets or common EU-membership.

In Table 8, we provide two extensions to our main PPML regressions from Table 3. First, we proceed similarly to our tetrad-regressions and add an interaction term between our post-withdrawal dummy and a dummy taking the value of one for all years after 2008. This way, we split the treatment–period by half to see whether the effect is stronger or weaker directly after the withdrawal as compared to some years down the road. Next, we restrict our sample to EU countries only to focus on Estonia's and Ukraine's main trading partners.

The results from splitting the treatment–period are presented in Columns (1) and (2) of Table 8, where in Column (2) we look at the directional effects. For the pooled effect in Column (1), we see only a slight difference between periods. The coefficient for the baseline–treatment is statistically significant and points to a 35% reduction in average trade volumes between Estonia and Ukraine. The point-estimate of the interaction with the dummy indicating the later period suggests an additional decrease in trade of around 23%. However, this estimate is only significant at the 10 percent level and hence should be interpreted with a grain of salt. The directional effects in Column (2) suggest no significant difference in the directional ef-

fects over time. For each direction, either the baseline or the interaction is significant and the respective other point estimate insignificant, supporting the findings expected from Column (1) that the effect persists and does neither increase nor fade out over time. Again, as above in Table 3, an F-Test suggests that trade decreased similarly in both directions between Ukraine and Estonia.
Looking at Columns (3)–(6), we see that the negative effect of FTA withdrawal on trade between Estonia and Ukraine yields a higher point estimate than in the full sample. However, the point estimates for both directional and un-directional effects are not significantly different from those estimates obtained for the full sample. What is new, though, is that the interaction with the post-2008 dummy variable is now throughout statistically significant at the one-percent level. Hence, when comparing trade between Estonia and Ukraine only to EU-countries, the magnitude of the withdrawal-effect seems to be increasing over time. Remarkably, the point-estimates of the interaction-terms are of a similar size as the baseline estimates. This hints at an accumulation of the trade destruction between Estonia and Ukraine over time, with an additional medium-term effect that reduces trade at a similar size as the direct effect after the withdrawal. We can conclude, hence, that the negative effect of the FTA withdrawal with high likelihood did not abate over the years but, if at all, shows a tendency to become more severe over time.

6. Conclusion

This paper studies the trade effects of FTA withdrawal by drawing on a unique natural experiment. When Estonia joined the European Union, it needed to withdraw from its FTA with Ukraine as part of the *acquis* of joining the Union. Since Estonia-Ukraine trade was relatively small, compared to the potential gains from EU membership, this was a trade-off worth accepting for Estonia. Hence, "Uxit" is plausibly exogenous and not related to any Estonia-Ukraine specific shocks.

We estimate the standard gravity model in a tetradic framework, with Latvia used as a counterfactual for Estonia. For Ukraine, we consider both Russia and the EFTA bloc as possible counterfactuals; the choice of counterfactuals is based on the economic and historical fundamentals of the region. The results point to a significant negative effect of FTA withdrawal on trade. Further, our estimate of the trade-destruction effect from FTA withdrawal falls in the same order of magnitude as the trade-creation effects of FTA formation found in the literature. This suggests that the withdrawal of an FTA undoes all trade-promoting effects the FTA brought in the first place, and that no "trading capital" is left behind. Further, we find no evidence of a delayed effect: the trade reduction follows immediately after the withdrawal. Rather, we find some suggestive evidence in some PPML specifications that the negative effect accumulates over time. However, we find clear evidence of an anticipation effect, as we find that trade between Estonia and Ukraine already went down significantly after the withdrawal was announced, but not yet in place. Our results are robust across different settings, first relaxing the choice of the reference-exporter for Ukraine by considering both potential tetrads in a panel setting as well as when we abandon the tetrad-setting completely and estimate a PPML model using data for 147 countries. Finally, we find based on a General Equilibrium model that both Ukraine and Estonia suffered significant welfare losses from the FTA withdrawal. Whereas the negative welfare effects for Estonia were more than offset by the benefits of joining the European Union, the welfare losses for Ukraine are further aggravated due to trade diversion effects.

Our central finding, that FTA withdrawal reduces bilateral trade by more than 20% and reverses all trade-effects gained by signing the FTA, is of huge importance for policy makers. The past years have seen significant efforts to reverse globalization and undo trade agreements that have been signed in the past, with Brexit being a case in point. Our findings provide first "real world" empirical evidence for an effect already hinted at in the literature: that the reversal of international economic cooperation significantly reduces trade between countries and, following standard trade-theory, may hence hurt the welfare of consumers.

References

Adam, C. and D. Cobham (2007). Modelling multilateral trade resistance in a gravity model with exchange rate regimes. *CDMA Conference Paper Series 0702.*

Anderson, J. E. and E. Van Wincoop (2003). Gravity with gravitas: A solution to the border puzzle. *The American Economic Review 93*(1), 170–192.

Anderson, J. E. and Y. V. Yotov (2016). Terms of trade and global efficiency effects of free trade agreements, 1990–2002. *Journal of International Economics 99*, 279–298.

Baier, S. and J. Bergstrand (2007). Do free trade agreements actually increase members' international trade? *Journal of International Economics 71*(1), 72–95.

Baier, S. L., J. H. Bergstrand, and J. P. Bruno (2019). Putting Canada in the penalty box: Trade and welfare effects of eliminating the North American Free Trade Agreement. *The World Economy 42*(12), 3488–3514.

Baier, S. L., J. H. Bergstrand, and M. Feng (2014). Economic integration agreements and the margins of international trade. *Journal of International Economics 93*(2), 339–350.

Baier, S. L., Y. V. Yotov, and T. Zylkin (2019). On the widely differing effects of free trade agreements: Lessons from twenty years of trade integration. *Journal of International Economics 116*, 206–226.

Berlingieri, G., H. Breinlich, and S. Dhingra (2018). The impact of trade agreements on consumer welfare—Evidence from the EU common external trade policy. *Journal of the European Economic Association 16*(6), 1881–1928.

Broda, C. and D. E. Weinstein (2006). Globalization and the gains from variety. *The Quarterly Journal of Economics 121*(2), 541–585.

Dhingra, S., H. Huang, G. Ottaviano, J. Paulo Pessoa, T. Sampson, and J. Van Reenen (2017). The costs and benefits of leaving the EU: trade effects. *Economic Policy 32*(92), 651–705.

Eicher, T. S., C. Henn, and C. Papageorgiou (2012). Trade creation and diversion revisited: Accounting for model uncertainty and natural trading partner effects. *Journal of Applied Econometrics 27*(2), 296–321.

Fajgelbaum, P. D., P. K. Goldberg, P. J. Kennedy, and A. K. Khandelwal (2020). The return to protectionism. *The Quarterly Journal of Economics 135*(1), 1–55.

Felbermayr, G., J. K. Gröschl, and I. Heiland (2018). Undoing Europe in a new quantitative trade model. *Ifo working paper*.

Felbermayr, G., J. K. Gröschl, and M. Steininger (2018). Quantifying Brexit: From ex post to ex ante using structural gravity. *CESifo Working Paper Series 7357.*

Feldmann, M. and R. Sally (2002). From the Soviet Union to the European Union: Estonian trade policy, 1991–2000. *The World Economy 25*(1), 79–106.

Gateva, E. (2016). European Union enlargement conditionality. *Palgrave Studies in European Union Politics ISBN: 978-1-137-48243-3.*

Head, K. and T. Mayer (2014). Gravity equations: Workhorse, toolkit, and cookbook. In *Handbook of International Economics*, pp. 131–195. Elsevier.

Head, K., T. Mayer, and J. Ries (2010). The erosion of colonial trade linkages after independence. *Journal of International Economics 81*(1), 1–14.

Khandelwal, A. K., P. K. Schott, and S.-J. Wei (2013). Trade liberalization and embedded institutional reform: Evidence from Chinese exporters. *American Economic Review 103*(6), 2169–95.

Mayer, T., V. Vicard, and S. Zignago (2019). The cost of non-Europe, revisited. *Economic Policy 34*(98), 145–199.

Moser, C. and A. K. Rose (2014). Who benefits from regional trade agreements? The view from the stock market. *European Economic Review 68*(C), 31–47.

Norkus, Z. (2007). Why did Estonia perform best? The North–South gap in the post-socialist economic transition of the Baltic states. *Journal of Baltic Studies 38*(1), 21–42.

Oberhofer, H. and M. Pfaffermayr (2017). Estimating the trade and welfare effects of Brexit. A panel data structural gravity model. *CESifo Working Paper No. 6828.*

Pain, N. and G. Young (2004). The macroeconomic impact of UK withdrawal from the EU. *Economic Modelling 21*(3), 387–408.

Rodrik, D. (2018). What do trade agreements really do? *Journal of Economic Perspectives 32*(2), 73–90.

Romalis, J. (2007). NAFTA's and CUSFTA's impact on international trade. *The Review of Economics and Statistics 89*(3), 416–435.

Santos Silva, J. and S. Tenreyro (2006). The log of gravity. *The Review of Economics and statistics 88*(4), 641–658.

Santos Silva, J. and S. Tenreyro (2011). Further simulation evidence on the performance of the poisson pseudo-maximum likelihood estimator. *Economics Letters 112*(2), 220–222.

World Bank (1995). The agreement between Estonia and Ukraine. available under: https:://wits.worldbank.org/GPTAD/PDF/archive/Estonia-Ukraine.pdf, World Bank.

World Bank (2014). Evaluation of the EU-Turkey customs union. Report No. 85830-TR, World Bank.

Appendix

A. Descriptives

TABLE A1
Descriptives Comext Dataset

		Monthly Import Volumes			
Importer & Partner	N	Mean	St. Dev.	Min	Max
EE all	19,943	7.133	20.919	0.00000	259.859
LV all	17,852	6.783	19.767	0.000	284.512
RU all	768	2,097.967	3,824.823	14.704	14,869.330
UA all	768	124.191	217.338	1.049	921.533
EE-UA	192	7.753	5.752	1.049	38.591
EE-RU	192	62.547	32.580	16.633	224.798
LV-UA	192	8.847	4.259	1.737	22.232
LV-RU	192	61.300	29.465	14.704	140.258

Notes: Descriptive Statistics for Comext-Data used in OLS Tetrad Regressions. Numbers represent total monthly import values in Euro and are displayed by reporting Country (Importer). The first four rows display aggregate imports for each importer in the sample, the last four rows show import flows for each importer and exporter in the sample, respectively.

TABLE A2
Descriptives, PPML Main Sample

Statistic	N	Mean	St. Dev.	Min	Max
trade	372,463	457,329,278.000	4,485,642,823	0	444,386,000,000
ftawithdrawn	714,210	0.00003	0.005	0	1
nr_pta	714,210	0.108	0.310	0	1
fta	714,210	0.037	0.189	0	1
cu	714,210	0.008	0.091	0	1
cm	714,210	0.010	0.101	0	1
eu	714,210	0.012	0.110	0	1
year	714,210	2002	–	1992	2012

Notes: Descriptive Statistics for Data used in Main PPML Regressions. Trade values are taken from Comtrade dataset and represent total yearly dyadic trade values in Euro.

TABLE A3
Descriptives, PPML Welfare Sample

Statistic	N	Mean	St. Dev.	Min	Max
trade	21,672	2,730,160,118	10,680,074,924	0	315,607,613,440
ftawithdrawn	21,672	0.001	0.024	0	1
nr_pta	21,672	0.120	0.325	0	1
fta	21,672	0.200	0.400	0	1
cu	21,672	0.019	0.137	0	1
cm	21,672	0.140	0.347	0	1
eu	21,672	0.159	0.366	0	1
year	21,672	2003.5	–	1998	2009

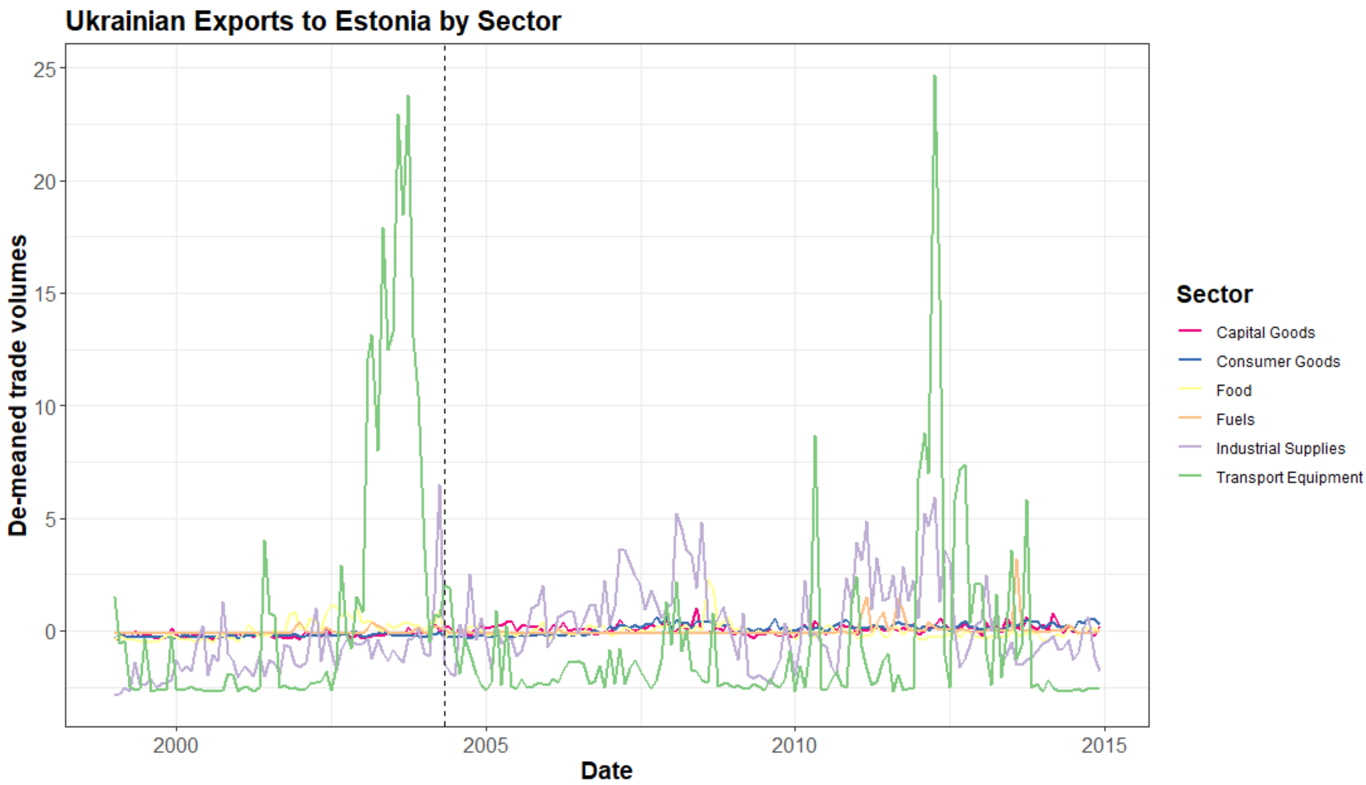
Notes: Descriptive Statistics for Data used in Main PPML Regressions. Trade values are taken from BACI dataset and represent total yearly dyadic manufacturing trade values in Euro.

B. Additional Figures

FIGURE B1
Construction of Dependent Variable

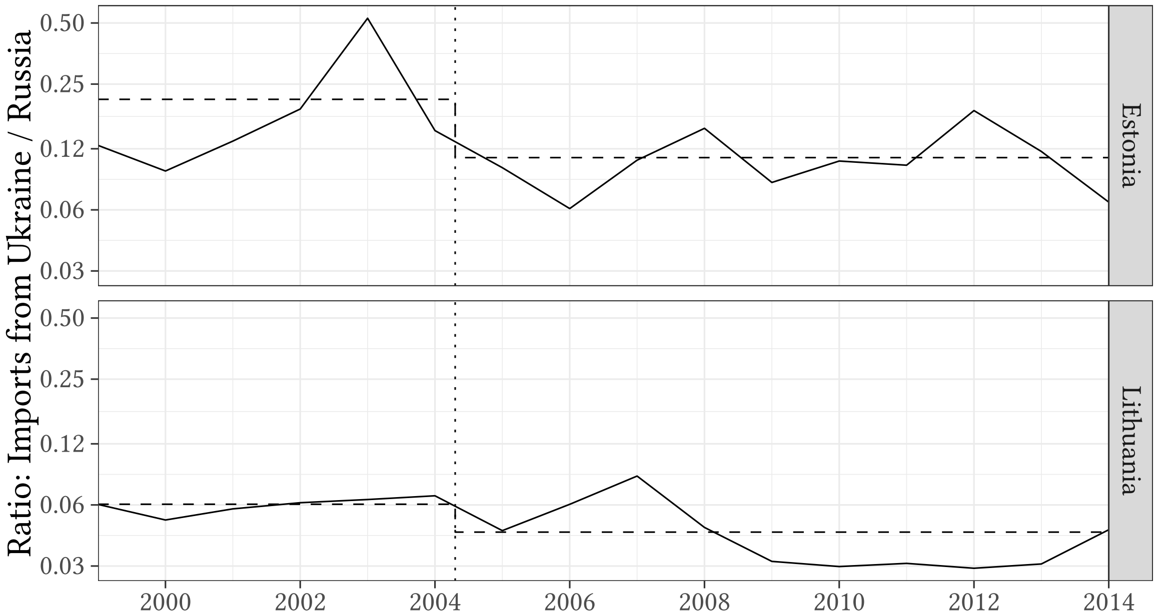
Notes: Graph illustrates the construction of the dependent variable used in the Tetrad regressions. The upper panel illustrates the case for Russia as reference exporter for Ukraine, the lower panel for EFTA as reference exporter. Green and blue lines indicate, for the importers Estonia (blue) and Latvia (green), the relative import penetrations of Ukraine vs. its respective reference Exporter. The red line illustrates the difference of these two import penetrations, which in effect constitutes our dependent variable. The vertical dashed line indicates the timing of the A8 EU Accession in May 2004. Calculated based on trade values in Euro provided by the Comext database.

FIGURE B2
Imports by Sector



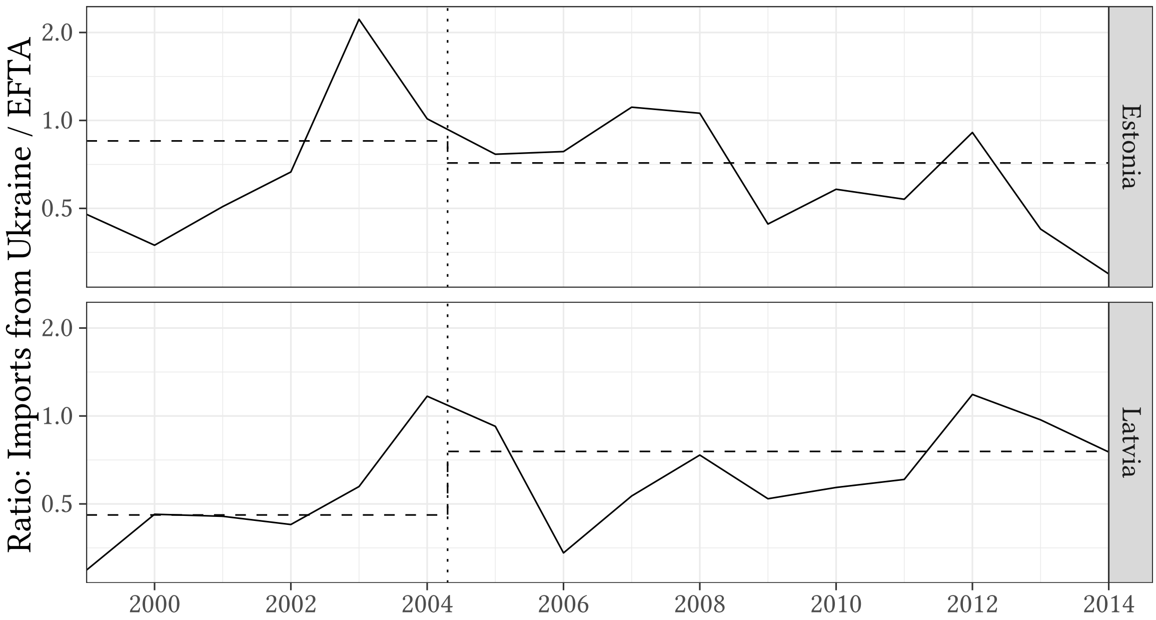
Notes: Graph illustrates Estonia's Imports from Ukraine disaggregated by sector. Calculated based on trade values in Euro provided by the Comext database. Line show monthly deviations from mean trade value for each product. The most volatile sector, indicated in green, is trade in transport equipment. The vertical dashed line indicates the timing of the A8 EU Accession in May 2004.

FIGURE B3
Trends in Import Penetration by Ukraine, Relative to Russia



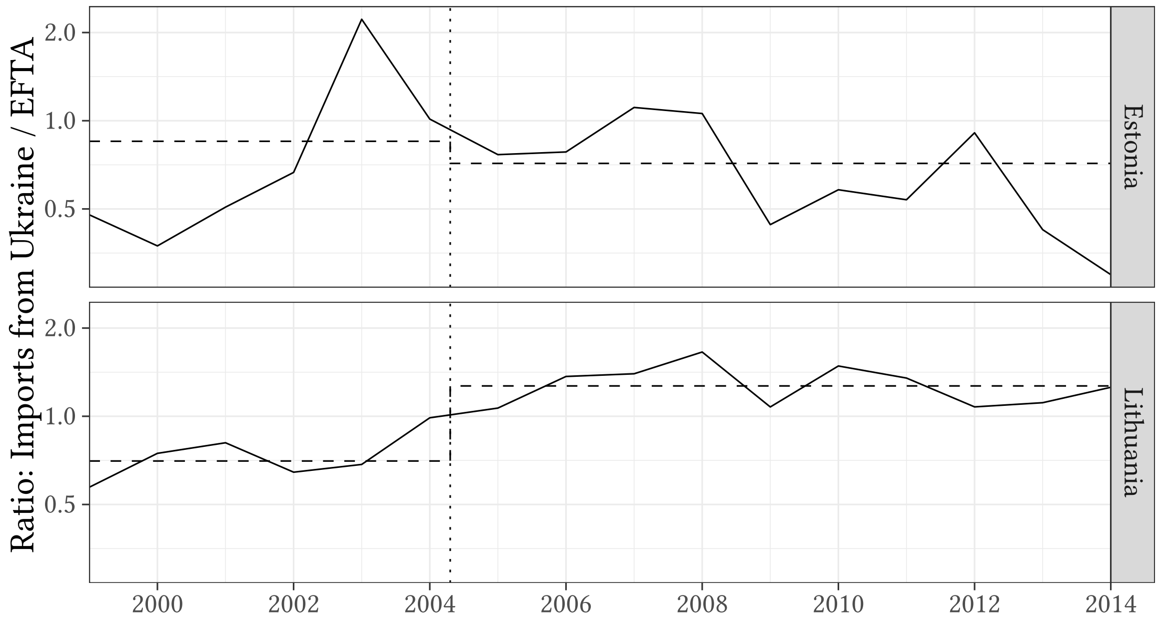
Notes: Graph illustrates the relative import penetration in Estonia and Lithuania, respectively, of imports from Ukraine vs. Russia. The vertical dashed line indicates the timing of the A8 EU Accession in May 2004. Calculated based on trade values in Euro provided by the Comext database.

FIGURE B4
Trends in Import Penetration by Ukraine, Relative to EFTA



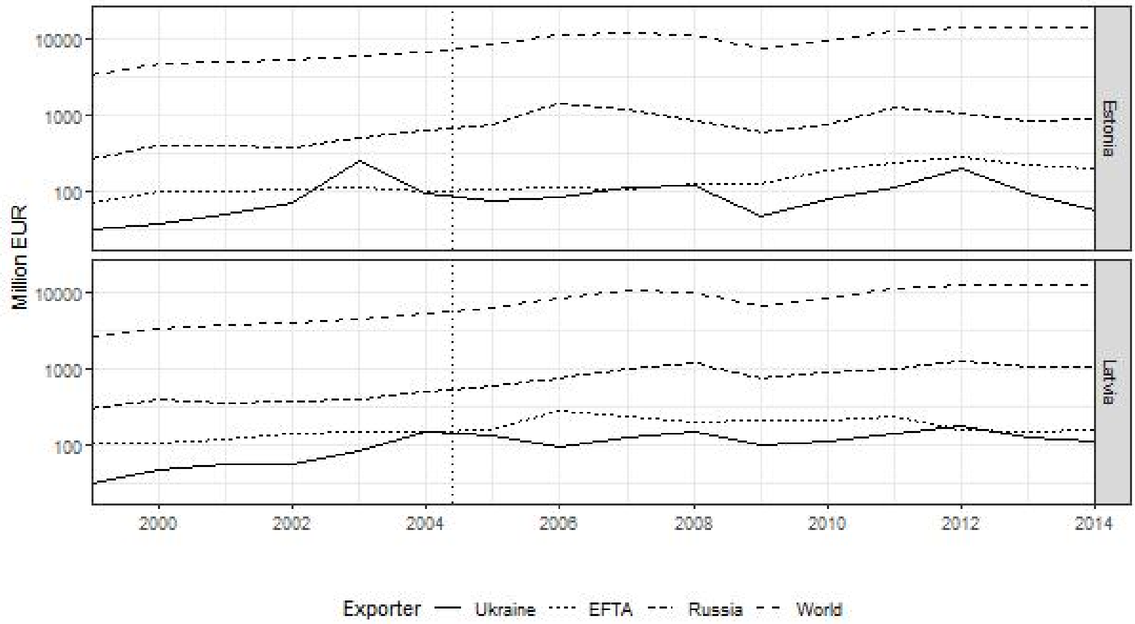
Notes: Graph illustrates the relative import penetration in Estonia and Latvia, respectively, of imports from Ukraine vs. the EFTA countries. The vertical dashed line indicates the timing of the A8 EU Accession in May 2004. Calculated based on trade values in Euro provided by the Comext database.

FIGURE B5
Trends in Import Penetration by Ukraine, Relative to EFTA



Notes: Graph illustrates the relative import penetration in Estonia and Lithuania, respectively, of imports from Ukraine vs. the EFTA countries. The vertical dashed line indicates the timing of the A8 EU Accession in May 2004. Calculated based on trade values in Euro provided by the Comext database.

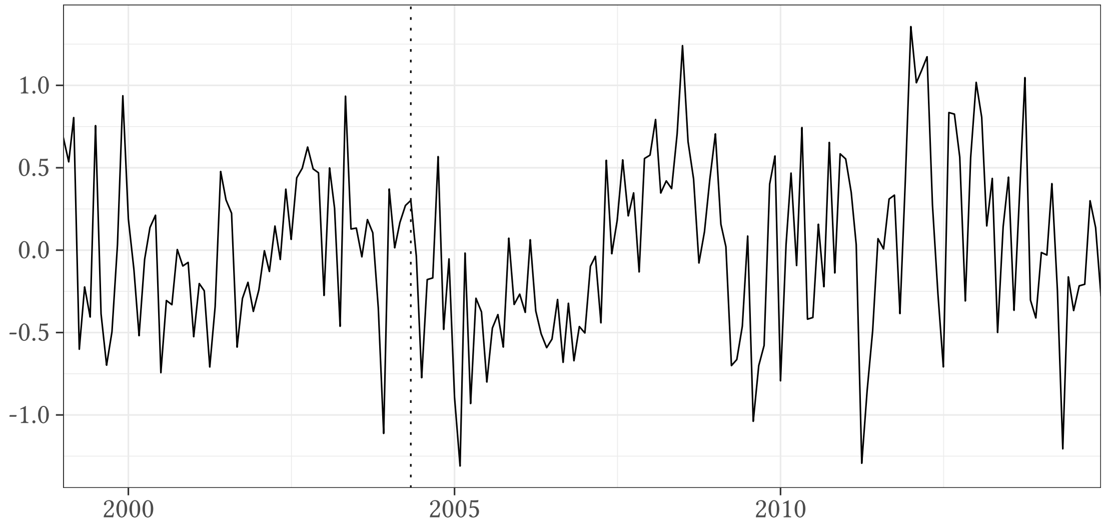
FIGURE B6
Baltic Trade Trends



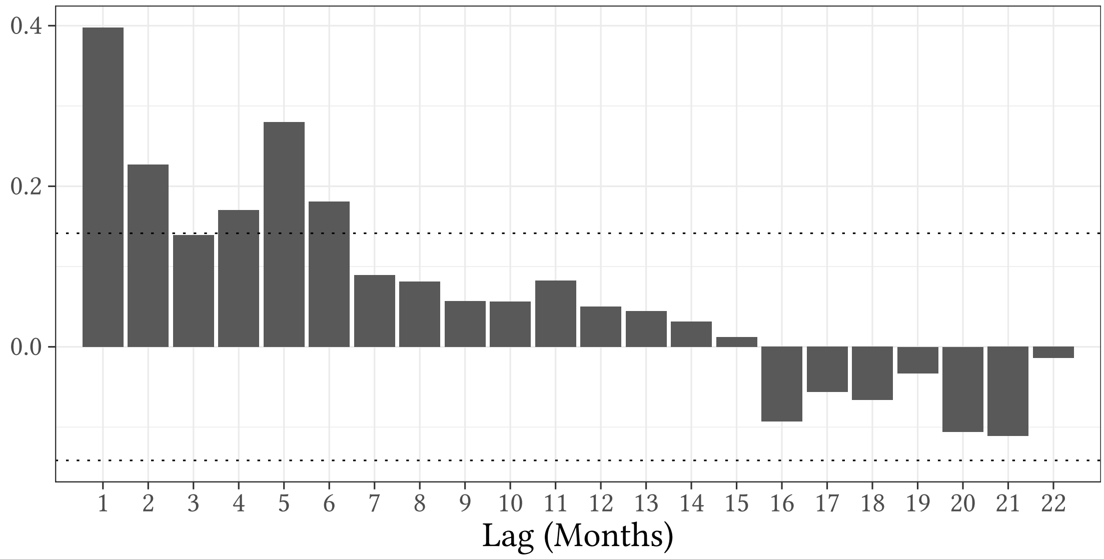
Notes: Graph shows the monthly import flows to Estonia and Latvia, respectively, from Ukraine, Russia, the EFTA countries and the world. The vertical dashed line indicates the timing of the A8 EU Accession in May 2004. Calculated based on trade values in Euro provided by the Comext database.

FIGURE B7
Regression Diagnostics, Model (2)

(a) Regression residuals



(b) Residual Autocorrelation Function

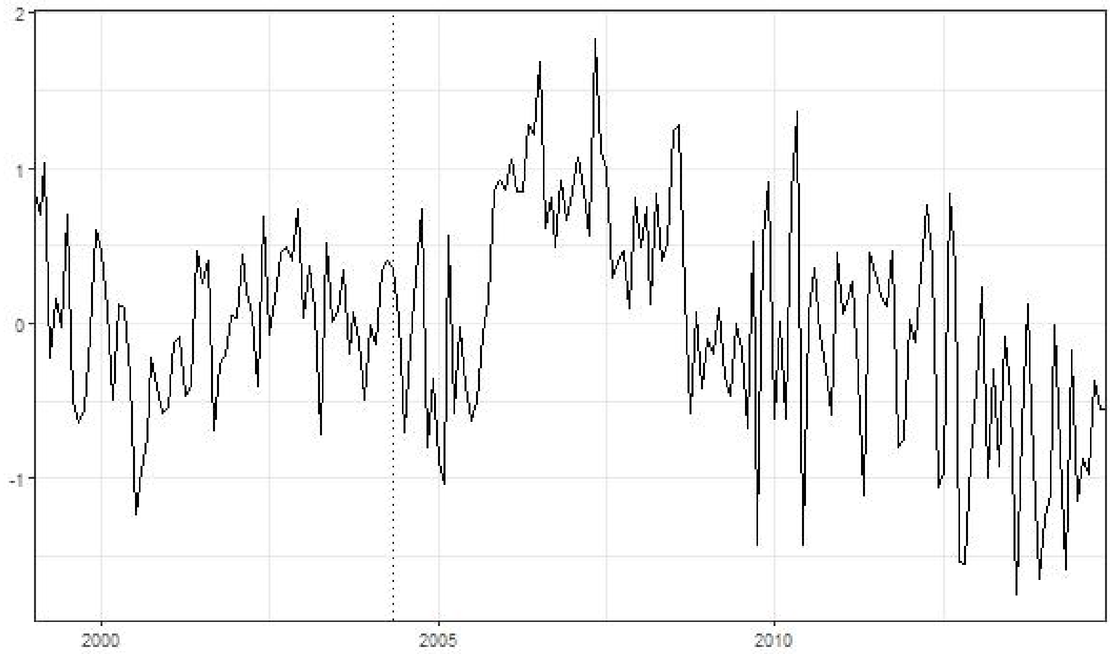


Notes: Graphs show the residuals from the OLS regression in Table 1, Column (2). Panel (a) plots the monthly residuals, Panel (b) shows the average autocorrelation between adjacent months. The vertical dashed line in Panel (a) indicates the timing of the A8 EU Accession in May 2004.

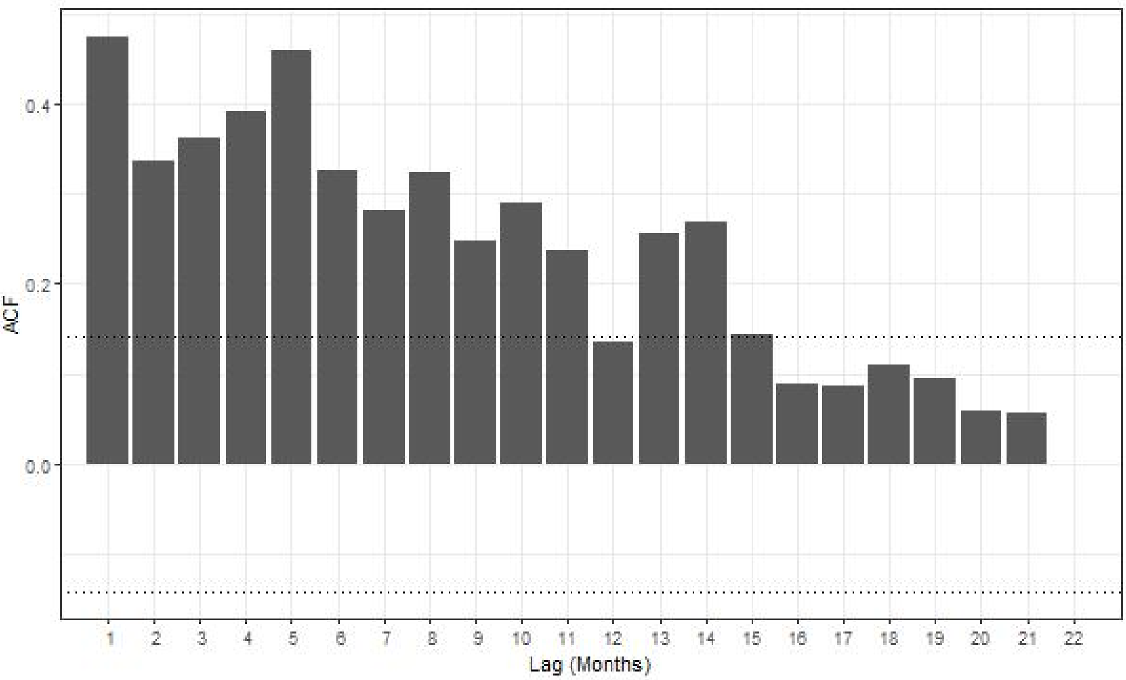
FIGURE B8

Regression Diagnostics, Model (5)

(a) Regression residuals

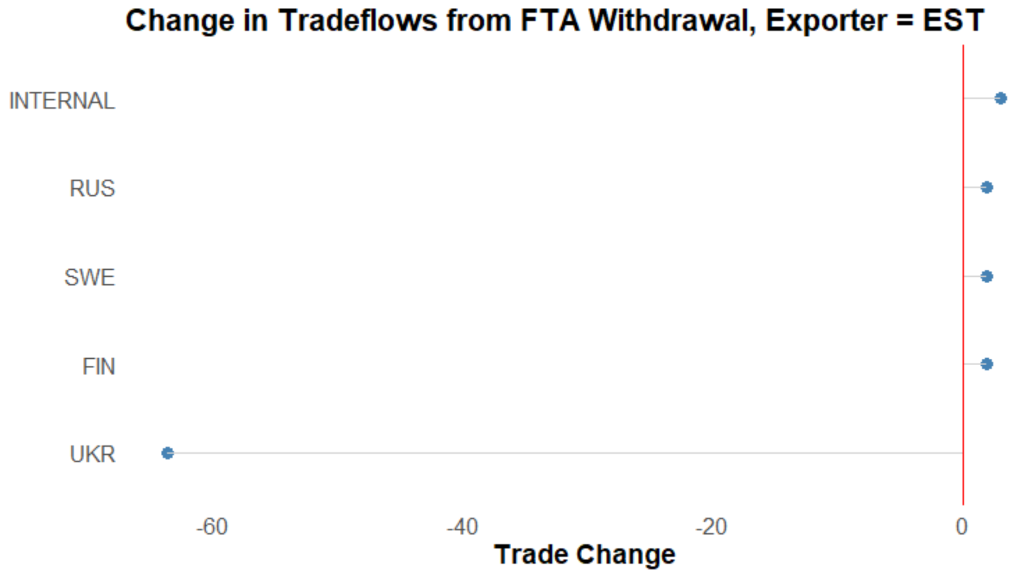


(b) Residual Autocorrelation Function

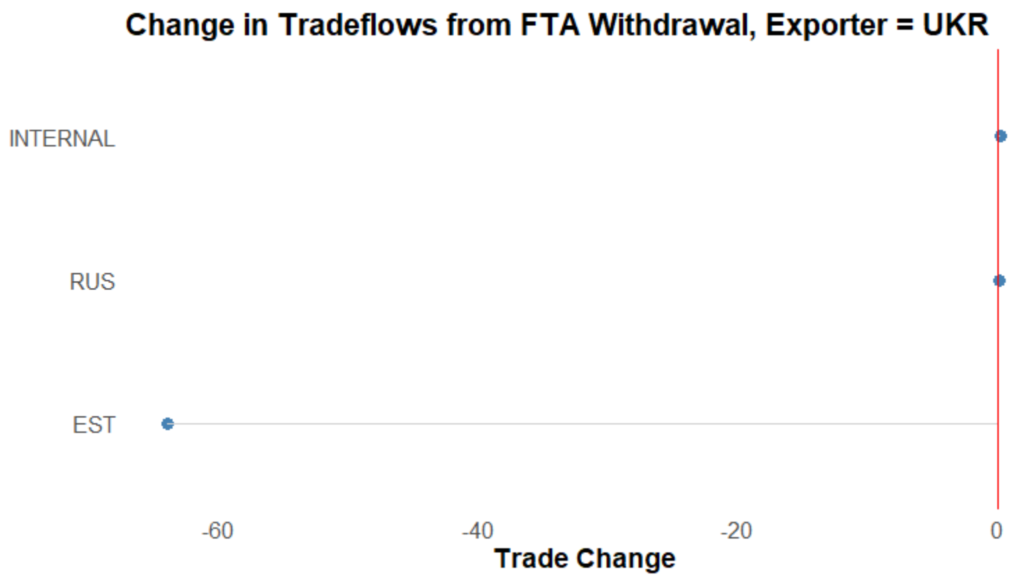


Notes: Graphs show the residuals from the OLS regression in Table 1, Column (5). Panel (a) plots the monthly residuals, Panel (b) shows the average autocorrelation between adjacent months. The vertical dashed line in Panel (a) indicates the timing of the A8 EU Accession in May 2004.

FIGURE B9
Trade Changes from FTA Withdrawal



(a) Estonia



(b) Ukraine

Notes: Graph shows the estimated changes in export flows Estonia and Ukraine, respectively, to various trading partners in percent. "INTERNAL" refers to a change in internal trade. The vertical red line indicates an estimated change of zero percent. Calculated based on General Equilibrium estimations.

C. Additional Tables

TABLE C1
Cross Correlations of Reference Countries

	EST	LVA	RUS	UKR
EST	1.00	0.94	0.7	0.76
LVA		1.00	0.75	0.78
RUS			1.00	0.9
UKR				1.00

(a) Change in log GDP

	EST	LVA	RUS	UKR
EST	1.00	0.94	0.68	0.78
LVA		1.00	0.75	0.8
RUS			1.00	0.91
UKR				1.00

(b) Change in log GDP per Capita

	EST	LVA	RUS	UKR
EST	1.00	0.92	0.82	0.84
LVA		1.00	0.9	0.86
RUS			1.00	0.96
UKR				1.00

(c) Change in log Total Imports

	EST	LVA	RUS	UKR
EST	1.00	0.8	0.93	0.87
LVA		1.00	0.82	0.89
RUS			1.00	0.91
UKR				1.00

(d) Change in log Total Exports

Source: GDP data are sourced from the IMF *World Economic Outlook*, October 2016 release. The series are NGDPD (GDP) and NGDPDPC (GDP per capita). World Import data are sourced from COMTRADE. All tables are based on annual data for the period 1999–2014.

TABLE C2
Cross Correlations of Reference Countries. Reference–Importer = Lithuania

	EST	LIT	RUS	UKR
EST	1.00	0.96	0.7	0.76
LIT		1.00	0.77	0.79
RUS			1.00	0.9
UKR				1.00

(a) Change in log GDP

	EST	LIT	RUS	UKR
EST	1.00	0.95	0.68	0.78
LIT		1.00	0.78	0.82
RUS			1.00	0.91
UKR				1.00

(b) Change in log GDP per Capita

	EST	LIT	RUS	UKR
EST	1.00	0.89	0.82	0.84
LIT		1.00	0.94	0.92
RUS			1.00	0.96
UKR				1.00

(c) Change in log Total Imports

	EST	LIT	RUS	UKR
EST	1.00	0.91	0.93	0.87
LIT		1.00	0.93	0.94
RUS			1.00	0.91
UKR				1.00

(d) Change in log Total Exports

Source: GDP data are sourced from the IMF *World Economic Outlook*, October 2016 release. The series are NGDPD (GDP) and NGDPDPC (GDP per capita). World Import data are sourced from COMTRADE. All tables are based on annual data for the period 1999–2014.

TABLE C3
Tetradic Regression Results. Reference–Importer = Lithuania

	(1)	(2)	(3)	(4)	(5)	(6)
ftaWithdrawn	-0.124	0.127	0.131	-0.658^{***}	-0.360^{***}	-0.388^{***}
	(0.103)	(0.100)	(0.107)	(0.111)	(0.101)	(0.108)
Reference Exporter	RU	RU	RU	EFTA	EFTA	EFTA
Controls						
Years '03, '04		X	X		X	X
Months in '03, '04			X			X
Durbin-Watson	0.00	0.00	0.00	0.00	0.00	0.00
BreuschGodfrey	0.00	0.00	0.00	0.00	0.00	0.00
Breusch-Pagan	0.89	0.01	0.79	0.03	0.34	1
Jarque-Bera	0.11	0.35	0.08	0.1	0.08	0.01
Rainbow	0.01	0.04	0.17	0.05	0.33	0.62
Observations	192	192	192	192	192	192
R^2	0.008	0.209	0.250	0.157	0.393	0.431

Notes: Newey-West HAC standard errors in parentheses. $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$. Results from OLS regressions with the logarithm of the ratio of relative import penetrations as dependent variable. The unit of observation is the Tetrad-month from January 1999 to December 2014. The main explanatory variable is a dummy variable indicating months after May 2004. In the year controls specification, indicators are added for the year being 2003 or 2004. In the month controls specifications, additional indicators are added for each month from February to December in the years 2003 and 2004. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text.

TABLE C4
Tetradic Regression Adding 2002 Dummies

	(1)	(2)	(3)	(4)	(5)	(6)
ftaWithdrawn	-0.463^{***}	-0.212^{**}	-0.183^{*}	-0.502^{***}	-0.192	-0.163
	(0.089)	(0.094)	(0.103)	(0.114)	(0.125)	(0.140)
Reference Exporter	RU	RU	RU	EFTA	EFTA	EFTA
Controls						
Years ’02, ’03, ’04		X	X		X	X
Months in ’02, ’03, ’04			X			X
Durbin-Watson	0.00	0.00	0.00	0.00	0.00	0.00
Breusch-Pagan	0.71	0.06	0.97	0.13	0	0.83
Jarque-Bera	0.13	0.69	0.49	0.32	0.69	0.67
Observations	192	192	192	192	192	192
R^2	0.124	0.311	0.385	0.093	0.231	0.275

Results from OLS regressions with the logarithm of the ratio of relative import penetrations as dependent variable. The unit of observation is the Tetrad-month from January 1999 to December 2014. The main explanatory variable is a dummy variable indicating months after May 2004. In the year controls specification, indicators are added for the year being 2002, 2003 or 2004. In the month controls specifications, additional indicators are added for each month from February to December in the years 2002, 2003 and 2004. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text.

TABLE C5
Tetradic Regression Results including Calender Month-Fixed Effects

	(1)	(2)	(3)	(4)	(5)	(6)
ftaWithdrawn	-0.459^{***}	-0.274^{***}	-0.261^{***}	-0.495^{***}	-0.264^{**}	-0.256^{**}
	(0.090)	(0.086)	(0.091)	(0.116)	(0.116)	(0.126)
Reference Exporter	RU	RU	RU	EFTA	EFTA	EFTA
Controls						
Years ’03, ’04		X	X		X	X
Months in ’03, ’04			X			X
Controls		y.03, y.04	ym.03.01−ym.04.12		y.03, y.04	ym.03.01−ym.04.12
Calender Month FE	YES	YES	YES	YES	YES	YES
Durbin-Watson	0.00	0.00	0.00	0.00	0.00	0.00
Breusch-Pagan	0.47	0.13	0.69	0.15	0	0.28
Jarque-Bera	0.12	0.86	0.67	0.25	0.51	0.61
Observations	192	192	192	192	192	192
R^2	0.163	0.340	0.402	0.109	0.238	0.271

Notes: Newey-West HAC standard errors in parentheses. *p<0.1; **p<0.05; ***p<0.01. In the year controls specification, indicators are added for the year being 2003 or 2004. In the month controls specifications, additional indicators are added for each month from February to December in the years 2003 and 2004. All models include calendar month-fixed effects. For diagnostic tests, p-values are always quoted. Own calculation based on dataset described in the text.

TABLE C6
PPML Results

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sample:	After 1997	All	All	All	All	All	All
FTA Withdrawn							
Both	-0.672***	-0.551***				-0.230	
	(0.149)	(0.134)				(0.209)	
Exporter = UA			-0.611***				-0.231
			(0.149)				(0.271)
Exporter = EE			-0.469**				-0.245
			(0.210)				(0.303)
FTA in tact							
Both				0.583***		0.362**	
				(0.127)		(0.156)	
Exporter = UA					0.656***		0.436**
					(0.129)		(0.197)
Exporter = EE					0.483**		0.245*
					(0.194)		(0.134)
PTA	0.0246	0.00704	0.00706	0.00703	0.00705	0.00704	0.00706
	(0.0373)	(0.0365)	(0.0365)	(0.0365)	(0.0365)	(0.0365)	(0.0365)
FTA	0.0584***	0.0551**	0.0551**	0.0551**	0.0551**	0.0551**	0.0551**
	(0.0226)	(0.0231)	(0.0231)	(0.0231)	(0.0231)	(0.0231)	(0.0231)
Currency Union	-0.0429	0.0771	0.0771	0.0771	0.0771	0.0771	0.0771
	(0.0922)	(0.0648)	(0.0648)	(0.0648)	(0.0648)	(0.0648)	(0.0648)
Common Market	0.0608**	0.0619**	0.0619**	0.0619**	0.0619**	0.0619**	0.0619**
	(0.0240)	(0.0242)	(0.0242)	(0.0242)	(0.0242)	(0.0242)	(0.0242)
EU	0.0200	0.0717	0.0717	0.0718	0.0718	0.0717	0.0717
	(0.0468)	(0.0528)	(0.0528)	(0.0528)	(0.0528)	(0.0528)	(0.0528)
Observations	303147	369735	369735	369735	369735	369735	369735

Notes: Standard errors clustered at dyad–level in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Results from PPML Regressions with trade volumes as dependent variable. The unit of observation is the dyad-year level for 179 countries and the years 1992 to 2012. The main explanatory variables are dummies that indicate either the conclusion or withdrawal of the FTA between Estonia and Ukraine. We alternate between non-directional (Columns 1, 2, 4, 6) and directional dummy variables (Columns 3, 5, 7). All regressions include exporter-year, importer-year as well as dyad Fixed Effects and control for the existence of trade agreements between countries, i.e. FTAs, PTAs, Currency Unions, Common Markets or common EU-membership.

TABLE C7
Extensions PPML

Sample	World		EU 27 only			
	(1)	(2)	(3)	(4)	(5)	(6)
$Withdrawn$	-0.427**		-0.677***		-0.516***	
	(0.169)		(0.110)		(0.119)	
$Withdrawn \times Year > 2008$	-0.263*				-0.333***	
	(0.154)				(0.0747)	
$Withdrawn_{Exporter=UA}$		-0.571***		-0.673***		-0.532***
		(0.141)		(0.149)		(0.162)
$Withdrawn_{Exporter=EE}$		-0.229		-0.682***		-0.498***
		(0.215)		(0.105)		(0.113)
$Withdrawn_{Exporter=UA} \times Year > 2008$		-0.0839				-0.286***
		(0.0724)				(0.0809)
$Withdrawn_{Exporter=EE} \times Year > 2008$		-0.503***				-0.396***
		(0.0871)				(0.104)
PTA	0.00704	0.00706	-0.425***	-0.425***	-0.425***	-0.425***
	(0.0365)	(0.0365)	(0.0897)	(0.0905)	(0.0898)	(0.0906)
FTA	0.0551**	0.0551**	0.0118	0.0118	0.0121	0.0121
	(0.0231)	(0.0231)	(0.0637)	(0.0637)	(0.0637)	(0.0637)
Currency Union	0.0771	0.0771	0.800***	0.800***	0.800***	0.800***
	(0.0648)	(0.0648)	(0.109)	(0.109)	(0.109)	(0.109)
Common Market	0.0619**	0.0619**	0.0108	0.0108	0.0110	0.0110
	(0.0242)	(0.0242)	(0.0361)	(0.0361)	(0.0361)	(0.0361)
EU	0.0717	0.0717	0.0182	0.0182	0.0183	0.0183
	(0.0528)	(0.0528)	(0.0620)	(0.0620)	(0.0620)	(0.0620)
Observations	369735	369735	15694	15694	15694	15694

Notes: Standard errors clustered at dyad–level in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Results from PPML Regressions with trade volumes as dependent variable. The unit of observation is the dyad-year level for 179 countries and the years 1992 to 2012. The main explanatory variables are dummies that indicate either the conclusion or withdrawal of the FTA between Estonia and Ukraine. We alternate between non-directional (Columns 1, 3, 5) and directional dummy variables (Columns 2, 4, 6). All regressions include exporter-year, importer-year as well as dyad Fixed Effects and control for the existence of trade agreements between countries, i.e. FTAs, PTAs, Currency Unions, Common Markets or common EU-membership.

TABLE C8
PPML Results, Welfare Dataset

	(1)	(2)	(3)	(4)
FTA Withdrawn				
Both	-0.837***			
	(0.200)			
Exporter = UA		-0.598***		
		(0.150)		
Exporter = EE		-1.105***		
		(0.0923)		
FTA in tact				
Both			0.837***	
			(0.200)	
Exporter = UA				0.598***
				(0.150)
Exporter = EE				1.105***
				(0.0923)
PTA	0.0636	0.0631	0.0636	0.0631
	(0.0655)	(0.0655)	(0.0655)	(0.0655)
FTA	0.0255	0.0255	0.0255	0.0255
	(0.0242)	(0.0242)	(0.0242)	(0.0242)
Currency Union	0.133	0.133	0.133	0.133
	(0.0824)	(0.0824)	(0.0824)	(0.0824)
Common Market	0.0791***	0.0791***	0.0791***	0.0791***
	(0.0253)	(0.0253)	(0.0253)	(0.0253)
EU	-0.0857	-0.0857	-0.0857	-0.0857
	(0.0528)	(0.0528)	(0.0528)	(0.0528)
Observations	21672	21672	21672	21672

Notes: Standard errors clustered at dyad–level in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Results from PPML Regressions with trade volumes as dependent variable. The unit of observation is the dyad-year level for 179 countries and the years 1992 to 2012. The main explanatory variables are dummies that indicate either the conclusion or withdrawal of the FTA between Estonia and Ukraine. We alternate between non-directional (Columns 1 and 3) and directional dummy variables (Columns 2 and 4). All regressions include exporter-year, importer-year as well as dyad Fixed Effects and control for the existence of trade agreements between countries, i.e. FTAs, PTAs, Currency Unions, Common Markets or common EU-membership.

D. Construction of GE Dataset

We require a symmetric dataset including internal trade flows for the estimation of welfare effects. Internal trade (internal consumption) is calculated as the difference between total production and exports. Production data is available only for manufacturing, hence, our welfare analysis is limited to manufacturing trade. To compute countries' internal trade flows, i.e. the volume of manufacturing production consumed internally and not shipped outside the country, we combine data from two sources. First, we use information on countries' total manufacturing production, provided by UN INDTSAT, and subtract the total amount of manufacturing exports based on the data provided in the World Bank's WITS dataset. Finally, we combine this measure of internal trade, coded as dyadic trade flows where importer and exporter are identical, with dyadic external trade flows of manufactured goods from the BACI dataset provided by CEPII. In the next step, we have to obtain a symmetric dataset, where all importers trade with all exporters, information on all countries' internal trade flows are included, and each exporter and importer appears every year. These requirements reduce our sample to the years 1999-2016, as well as to 56 importers and exporters. We keep years as well as importers/exporters in our sample whose observations amount to at least 90% of the observations in the year or importer/exporter, respectively, with most observations in the sample. We repeat this exercise both for stricter (e.g. 99%) and looser (e.g. 70%) restrictions, but our results remain unchanged. Finally, we fill any remaining missing trade flows between any two countries with zeroes. In total, we add 342 zeroes for international flows, which account for 0.642% of our overall sample. Note that looser restrictions in dropping sample countries/years result in more zero trade flows, whereas being stricter (e.g. 95%) would not have required adding any zeros. We also did not add any zero trade flows for internal trade observations, but dropped all countries with at least one year of missing internal trade flows. A final complication accrues due to negative internal trade flows, a problem that has already been noted by Baier, Yotov, and Zylkin (2019). Such negative internal trade flows often occur in rather small but open countries with a high export-to-GDP ratio and may result from inaccuracies in the reported data. We follow the convention in the literature and drop countries for which negative internal trade flows are reported for at least one year from our sample (these are Belgium, Hong Kong, Singapore, and Slovakia). Unfortunately, also Estonia, our country of interest, reported negative trade flows in the years 2000 and 2001. Therefore, we have to drop those two years from our regressions.