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Trade effects of standards harmonization in the EU: improved access for non-EU partners

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Abstract

In September 2008, the EU replaced national-level regulation on Maximum Residue Levels (MRLs) in pesticides with harmonized Community-wide regulation. Using data on pesticide MRLs for 53 trading partners over 2005-2014, we examine the effects of this harmonization on both intra- and extra-EU imports in an original empirical contribution to this literature. We also embed regulatory heterogeneity in the Melitz (2003) framework in a theoretical contribution. We find robust evidence for adverse effects of regulatory heterogeneity on intra-EU trade in the pre-harmonization period, which questions the implementation of the Cassis de Dijon principle. Our findings also suggest that the EU's MRL harmonization may have improved access for non-EU, including non-OECD, exporters to the Common Market.

Key words: *Standards, MRL Regulation, Harmonization, EU, Trade*

JEL classification: *F13, F14, I18*

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Anirudh Shingal* and Malte Ehrich†

1 Introduction

The continual decline of tariffs through successive rounds of multilateral trade negotiations has increased the relative importance of non-tariff measures (NTMs). Sanitary and Phytosanitary (SPS) standards and technical barriers to trade (TBT) are two such NTMs, which though imposed for legitimate reasons such as alleviating information asymmetries, mitigating consumption risks and promoting environmental sustainability, can also be instruments of disguised protectionism.

A commonly used standard in agricultural products restricts the maximum residue level (MRL) from pesticides. A pesticide residue is a tiny trace of pesticide that sometimes remains on the treated crop. A MRL is the maximum amount of residue legally permitted on food products. Once residues are demonstrated to be safe for consumption, MRLs are set by scientists, based on rigorous evaluation of each legally authorized pesticide. Countries choose the products they regulate, the pesticides they regulate for each product, as well as the MRL for a given product-pesticide pair.

The standards literature has studied the impact of MRL regulation on trade.[1] In this paper, we re-visit the effect of regulatory heterogeneity on bilateral trade using the near-natural experiment setting provided by the harmonization of MRL regulation within the EU, which has not yet been studied in this literature.

Regulation (EC) No 396/2005 contains a list of MRLs that came into effect in September 2008 and effectively repealed Member State MRL regulation from there onwards. Thus, before 1 September 2008, a mixed system was in place with harmonized Community MRLs for ca. 250 active substances and national MRLs for the remaining substances. After this date, harmonized MRLs became applicable for all active substances used in plant protection products that have the potential to enter the food chain.

We examine the effects of this MRL harmonization within the EU using the Homologa data[2] on pesticide MRLs over 2005-2014 for 53 exporting and importing countries (details in Section 5). In another original contribution, we incorporate both relative dyadic MRL restrictiveness and regulatory heterogeneity within the EU in the Melitz (2003) framework to

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1 The following section provides a review of the literature relevant to this paper. A more extensive review of the MRL-trade literature is provided in Shingal et al. (2017).

2 These data are obtained from Agrobase-Logigram, a private company that maintains Homologa, the Gobal Crop Protection Database.

motivate the empirical analysis.

Anecdotal evidence suggests that both differences in regulatory standards between the exporting and importing countries and diversity of standards within a Common Market like the EU can impose trade costs on exporting firms. For instance, in its report[3] on non-tariff barriers (NTBs) faced by Indian agricultural products, India's Agricultural and Processed Food Products Export Development Authority (APEDA) lists both (i) the requirement to meet more stringent standards in the EU and (ii) the lack of harmonization of products standards in EU Member States resulting in the need to approve products/production units by individual member countries, amongst NTBs faced by Indian agri-exporters in the EU.

The ease with which exporting country firms can meet stricter importing country standards depends on the level of regulation in the exporting country, the restrictiveness of importing country standards relative to the exporting country and on whether exporting country firms have a comparative advantage in meeting stricter regulation. It is important to remember that having to comply with more stringent regulation not only has a trade cost effect, but can also be associated with a demand-enhancing effect, especially if the exports are destined for markets where consumer preferences are more pro-food-safety.

At the same time, heterogeneity in regulatory standards in a common market such as the EU increases the fixed product adaptation costs that exporting firms must pay in order to access the common market, discouraging market entry and reducing the range of exported product varieties. This trade cost effect could even be prohibitive for poor countries, typically characterized by lack of access to information, technology, managerial capacity and finance, thus impeding the ability of firms to adapt production processes quickly and adequately to meet heterogeneous product standards in the EU, or to obtain testing and certification services required to demonstrate conformity.

Our findings suggest that MRL harmonization may have made it easier for the EU's non-EU-both OECD and developing country - partners to export to the Common Market though the effect is found to be stronger for OECD exporters. We also find robust evidence for regulatory heterogeneity within the EU in the pre-MRL-harmonized period to be associated with adverse effects on intra-EU trade, which questions the implementation of the Cassis de Dijon principle. The Cassis de Dijon principle is a cornerstone of the EU's internal market. It requires Member States to mutually recognise their national regulations in cases where there are no generally binding EU regulations. This means that goods produced and marketed in one EU Member State may be sold without further restrictions in all other Member States.

We also decompose our results by margins of trade using the Heckman (1979) two-step estimator following Helpman et al. (2008). These results suggest that the positive trade effects of the EU's MRL harmonization are primarily observed at the extensive margin, which is consistent with the findings in this literature (for instance see Shepherd, 2007).

The rest of the paper is structured as follows. The following section reviews some of the

[3] http://apeda.in/apedahindi/Databank/NTBs_March_08.pdf

relevant literature while Section 3 outlines the theoretical model motivating our empirical analysis. Section 4 describes the measures of regulatory heterogeneity that we use to examine trade effects of MRL harmonization in the EU. Section 5 presents the data while Section 6 discusses the empirical methodology and related issues. The results are discussed in Section 7 and Section 8 concludes.

2 Literature review

We are not the first to study the trade effect of product standards harmonization, either generally or in the context of the EU. Otsuki et al. (2001b,a) found the EU's harmonized aflatoxin standard to be associated with a 63% more decline in select African food exports relative to the standard set by Codex. However, their empirical strategy did not incorporate recent advancements in the estimation of structural gravity, leading to biased estimates.

Chen and Mattoo (2008) used a sample selection gravity model (following Helpman et al. 2008) to examine the impact of EU Harmonization Directives and Mutual Recognition Agree- ments on intra- and extra-EU trade. Baller (2007) adopted the same approach using data on both EU and ASEAN harmonization and mutual recognition agreements. Both studies found harmonization to boost trade among harmonizing countries, as well as imports from third countries. Shepherd (2007) examined the effect of the share of the EU's (CEN Euro- pean) standards in textile, clothing and footwear sectors identical to ISO standards on the variety of exports coming from non-EU countries into the EU and found a positive extensive margin effect.

Chen and Mattoo (2008) used the number of Harmonization Directives in a sector in an EU/EFTA Member State in their analysis while Baller (2007) constructed a dummy variable for sectors (telecom equipment and medical devices) harmonized between countries i and j. In contrast to these studies, we use a continuous dyadic measure of MRL restrictivness and that of regulatory heterogeneity within the EU (for details see Section 4), reflecting MRL harmonization in the EU post-2008 by construction and arguably, enabling a more direct identification of the treatment effect.

Achterbosch et al. (2009) studied the impact of differences in pesticide MRLs on Chilean fruits exports to the EU-15 over 1996-2007 and found a 5% reduction in the EU's regulatory tolerance levels for MRLs to lead to a 14.8% decline in export volumes, with grapes being twice as sensitive as the other fruits. They constructed an index of MRL restrictiveness by taking the difference in MRL regulation between countries in a trading dyad and normalizing it by the sum of the levels of MRL regulation in that dyad. However, the Achterbosch et al. (2009) index does not possess one of the desirable properties of restrictiveness indices (for instance see Li and Beghin, 2014), namely convexity in protectionism.

The paper closest to ours is De Frahan and Vancauteren (2006) that studied the trade effects of harmonization of food regulations in the EU on intra-EU trade in food products over 1990-2001 by considering harmonization initiatives in EC Directives. They found this harmonization to have a large and positive effect on import intensity both at the aggregate level and for individual food sectors. However like Baller (2007) and Chen and Mattoo (2008), the authors looked at trade flows associated with harmonization initiatives in EC Directives, which may not provide the cleaner identification that our continuous measures of regulatory heterogeneity provide.

Finally, consistent with the recent empirical trade literature (for instance see Baier et al. 2014), we also use three-way fixed effects to mitigate endogeneity-induced biases in the standards-trade relationship (for instance see Disdier et al. 2014, Shingal et al. 2017). This is different from the IV estimation approach used by Baller (2007), Shepherd (2007) and Chen and Mattoo (2008), which did not substantially change the empirical results in each respective case, relative to the OLS analysis.

3 Theory

3.1 The Melitz-model

The essential role of fixed costs for production and exports has been emphasized in the "New-trade-theory" as well as in the "New-new-trade theory". Whereas the former is mostly motivated to explain intra-industry trade by implementing product differentiation in a monopolistic competition framework, the latter relaxes the assumption of firm homogeneity by arguing that exporting firms have fundamentally different characteristics than non-exporting firms in terms of productivity, wages, production volumes, and profits (Mayer and Ottaviano, 2007; Colen et al. 2012). In what follows, we use elements of the heterogeneous firm of the Melitz model (Melitz, 2003) to demonstrate the effect of heterogeneity in product regulation on bilateral exports.

Melitz introduces firm heterogeneity via the productivity parameter ϕ. Firms need to pay sunk entry costs f_E to draw their productivity level from a cumulative Pareto distribution $G(\phi)$. This productivity level determines whether the firm exits the market, serves the domestic market only, or exports to foreign markets. Production requires fixed costs f_{ii} for serving the domestic market i and incorporates market access costs and fixed production costs. Export costs from country i to country j are denoted by f_{ij}. Hence, profits are given by Equation (1):

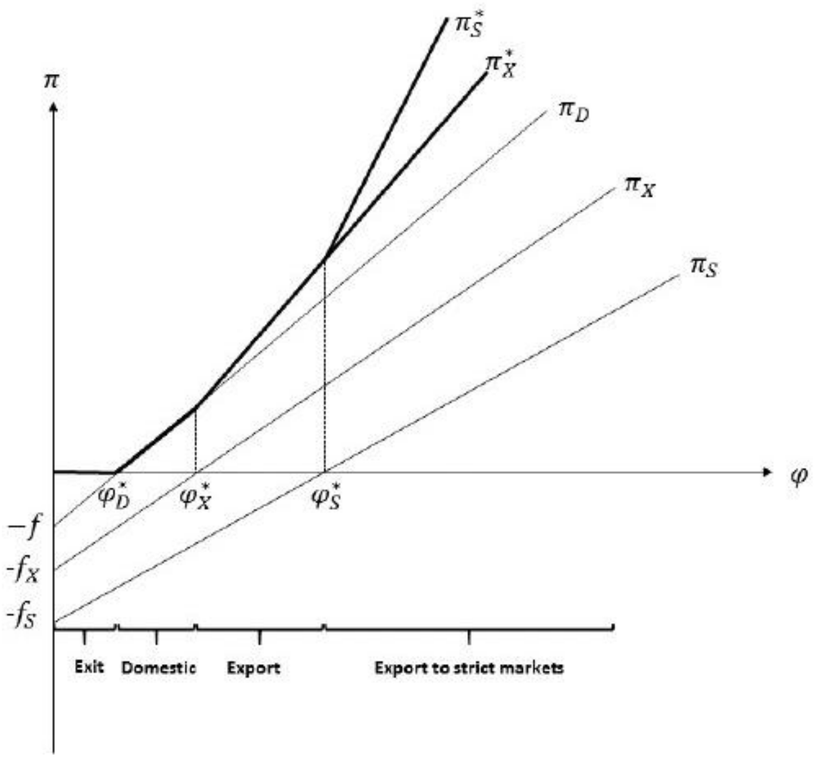
$$\pi_{ij}(\phi) = B_j \tau_{ij}{}^{1-\sigma} \phi_{ij}{}^{\sigma-1} - f_{ij} \tag{1}$$

where B_j is a demand parameter of the destination market j, τ represents iceberg-type trade costs, and ϕ is the drawn productivity parameter. σ denotes the constant elasticity of substitution which is assumed to be greater than one. Thus, the zero-profit cutoff condition for exporting from market i to market j yields the cutoff productivity level ϕ^*_{ij} at which profits are zero, see Equation (2).

$$\pi_{ij}(\phi^*_{ij}) = 0 \Longleftrightarrow B_j (\tau_{ij})^{1-\sigma} (\phi^*_{ij})^{\sigma-1} = f_{ij} \tag{2}$$

In equilibrium, higher fixed costs f_{ij} are associated with higher demand, lower trade costs, or higher productivity for $\sigma > 1$. Figure (1) depicts the relation between fixed costs and various cutoff productivity levels. If the drawn productivity level is below ϕ^*_D, the firm exits the market; if $\phi^*_D < \phi < \phi^*_X$, the firm produces for the domestic market only but does not export. Once the productivity level exceeds ϕ^*_X, the firm exports. Note that the slope of the corresponding profit curve π_X is smaller than for π_D due to variable trade costs. Profits for exporting firms are jointly determined by π_D and π_X and given by the bold curve π^*_X.

Figure 1: Profits, productivity, and standards in the Melitz framework



3.2 Heterogeneous firm theory and food standard heterogeneity

As argued in the introduction, compliance with food standards, especially in the presence of heterogeneity, requires additional fixed costs. Melitz already defines fixed costs broadly as "market access" costs. Therefore, the stricter (relative to domestic standards) and more heterogeneous food standards in a common market such as the EU, the larger are market access costs likely to be. We thus not only modify the Melitz-model to incorporate higher fixed costs due to the relative dyadic restrictiveness of standards but also those emanating from the heterogeneity of standards for the same pesticide-product combination across EU Member States. In particular, we add a firm-specific quality upgrading fixed cost term $f(q_i, \delta)$ where q_i is a firm-specific quality parameter for differentiated goods and δ denotes the heterogeneity of standards within the EU.

Moreover, we incorporate food standards in our theoretical framework not only via increased fixed costs of exporting on the supply side but also as a strategy to address preferences of modern consumers on the demand side and hence, as a form of quality upgrading (Ferguson, 2009). Thus, profits also increase in qi. The zero-profit cutoff condition (2) then changes to:

$$\underbrace{\max}_{q_i} \quad \left[q_i B_j \left(\tau_{ij}\right)^{1-\sigma} \varphi^{\sigma-1} - f(q_i, \delta) - f_{ij}\right] \tag{3}$$

We assume that the firm-specific quality upgrading fixed costs $f(q_i, \delta)$ are continuously differentiable. This implies that the conformance with a specific quality-level, and hence with specific standard-requirements, is not a binary decision. Instead, an optimal standard can be chosen from a broad continuum of standards. Following (Ferguson, 2009), we need to specify the functional form of $f(q_i, \delta)$. We assume that meeting relatively low levels of standards is a low hanging fruit. However, costs are expected to increase exponentially since it becomes increasingly difficult to meet high levels of standards. Thus, we assume that quality upgrading fixed costs are convex and increasing in qi. Hence, the partial derivative of $f(q_i, \delta)$ with respect to q_i increases in q_i. Moreover, we assume fixed costs to be higher for higher degrees of standard heterogeneity.

To be more explicit, we assume the following functional form of $f(q_i, \delta)$:

$$f(q_i, \delta) = q_i^{1/\theta}\delta \quad \text{with} \quad \theta \, \varepsilon [0, 1] \quad \text{and} \quad \delta \, \varepsilon [1, \infty) \tag{4}$$

The shape-parameter θ indicates the “ease” of quality upgrading (Ferguson, 2009, p.10). The larger is θ, the more easily can a firm address preferences of consumers that demand high-quality products; i.e. products that meet relatively strict standards. Hence, if firms are able to meet standards easily, they will benefit from a lower increase in associated costs which - eventually - meet higher levels of demand. Using this specific functional form of the costs of quality-upgrading, the optimal quality-level is then given by:

$$q^* = \left(\frac{\theta}{\delta} B_j \tau_{ij}^{1-\sigma} \varphi^{\sigma-1}\right)^{\frac{\theta}{1-\theta}} \tag{5}$$

Keeping B_j, τ_{ij}, and ϕ constant, higher values of θ increase the optimal level of quality. Thus, if a standard is particularly capable of addressing consumers’ preferences - i.e. high values of θ - producers can earn higher profits by investing in stricter - i.e. high-quality standards. In this scenario, we would expect standards to increase profits and therefore, increase trade flows at the aggregate level. In contrast, if a particular standard is less capable of addressing consumers’ preferences - θ is close to zero - the quality level remains low and producers are less likely to invest in the standard.

The other parameter of interest in our analysis is δ which represents the heterogeneity of standards in a common market such as the EU. In the case of full harmonization we assume δ to be equal to one. The more heterogeneous standards are the higher is δ. Equation (5) shows that an increase in δ is associated with lower levels of the optimal quality choice. In other words, a smaller set of firms will invest in standards of a given quality the larger δ becomes. This will result in lower trade.

4 Measures of regulatory heterogeneity

We proxy θ using an index of dyadic regulatory heterogeneity. Following Winchester et al. (2012), the index[4] is defined at the pesticide level as follows:

$$r_{ijpkt} = \frac{abs(MRL_{ipkt} - MRL_{jpkt})}{max(MRL_{pkt}) - min(MRL_{pkt})} \tag{6}$$

and at the product level, as follows:

$$R_{ijpt} = \frac{1}{K}\sum_{k=1}^{K} r_{ijpkt} \tag{7}$$

where MRL_{ipkt} is the maximum residue level of pesticide k allowed by the exporter i to remain on product p, MRL_{jpkt} is the maximum residue level of pesticide k allowed by the importer j to remain on product p[5] and $maxMRL_{pkt}$ and $minMRL_{pkt}$ are the maxima and minima of pesticides across all sample countries.

In constructing the index at the product level, we average over the number of pesticides as the number of pesticides regulated is found to vary by product. For instance, the US has 107 pesticide MRLs for apples but only seven pesticide MRLs for coconut (Li and Beghin, 2014). By averaging the sum of the heterogeneity index of each pesticide by the total number of pesticides, we make the indices invariant to regulation intensity *a la* Li and Beghin (2014). The use of the simple average thus avoids assigning higher values to certain products simply because a greater number of pesticides are commonly applied to those products.

The index, R_{ijpt}, thus measures the relative difference in MRL regulation between exporter i and importer j, regarding the maximum residue level of pesticide k, on average, allowed to remain on product p. The value of the index ranges from zero (the absence of dyadic regulatory heterogeneity) to one (maximum dissimilarity in regualtion between trading partners).

If we consider the index R_{ijpt}, a strong negative effect at the extensive margin would suggest that having dissimilar MRL regulation in a dyad is a fixed cost that producers have to overcome before being able to export towards a more stringent destination. The same effect at the intensive margin may suggest that the cost of complying with relatively stricter MRL regulation is variable and increases with the value of exports. Literature suggests that dyadic regulatory heterogeneity may affect both fixed and variable costs (for instance see Ferro et al. 2015, Xiong and Beghin, 2014, respectively).

[4] One advantage of using this index is that it fulfills all the desirable properties of heterogeneity indices viz. scale-invariance, convexity in protectionism, invariance to regulation intensity, monotonicity and having lower and upper bounds (for instance see Li and Beghin, 2014). Moreover, the index is truly dyadic, which is a requirement of our research objective.

[5] Thus, the MRL_{ipkt} and MRL_{jpkt} are non-negative variables, that are theoretically unbounded but bounded from above in practice.

On the other hand, the coefficient of R_{ijpt} could also be positive (for instance see Ishaq et al. 2016; Shingal et al. 2017; Li et al. 2017) which would indicate that stricter regulation has more of a demand-enhancing as opposed to a trade cost effect. This would be consistent with larger values of θ in our theoretical framework.

We proxy δ by the standard deviation ($\sigma^{EU}{}_{pt}$) of MRLs for each HS-6 digit product (averaged across pesticides used for that product) across the importing EU Member States. We expect this measure to be positive and significant over the pre-harmonization period (2004-2008) and zero thereafter (consistent with the harmonization of MRLs undertaken post-2008).

Empirically, we expect $\sigma^{EU}{}_{pt}$ to impact agri-exports to the EU from both EU and non-EU partners negatively, especially at the extensive margin, over the pre-harmonization period.

Finally, we would also like to point out a few cases in the construction of the heterogeneity indices. Not all countries set MRLs for the same pesticide/crop combination; it can therefore be the case that the importer country sets an MRL for a k, p pair for which the exporting country has not set a limit and we would therefore have to drop this observation as no comparison is possible. To minimize this from happening, and without imputing values arbitrarily, we resort to using default MRL values.[6] Some countries set default MRLs for any k, p combination that is not explicitly cited in their MRL regulation, such as the EU that sets an MRL of 0.01 mg/kg for any pesticide on any crop that is not listed in the European Commission Regulation No 396/2005.

Table 1 summarizes the pertinent default MRL cases. Thus, in cases where one of the partner countries was missing the MRL, we resort to the missing country's default value (if any) to compute the heterogeneity measures. In cases where there is no default MRL in place as well, we replace the missing MRL with the sample's highest MRL following Drogué et al. (2012), Ferro et al. (2015) and Shingal et al. (2017).

[6] Drogué and DeMaria, 2012, Xiong and Beghin (2014) and Shingal et al. (2017) also resort to the use of default values.

Table 1: Many countries use Codex MRLs as default values if national regulation is missing

Country	First default	Second default
Argentina	Codex	0.01
Australia	0.01	
Brazil	Codex	
Canada	0.01	
Chile	Codex	
China	Codex	
Egypt	Codex	
European Union	0.01	
India	Codex	
Israel	Codex	
Japan	0.01	
South Korea	Codex	
Malaysia	Codex	0.01
Mexico	Codex	
New Zealand	0.01	
Norway	0.01	
Russia	Codex	
Singapore	Codex	
South Africa	Codex	0.01
Switzerland	EU	0.01
Thailand	Codex	
Turkey	Codex	
Ukraine	Codex	
USA	0.01	
Vietnam	Codex	0.01

Note: Default MRL information from mrldatabase.com (US FDA) except otherwise stated.

In the empirical analysis that follows, we examine the sensitivity of our results both to the use of default MRL values and to replacements by the sample maxima by considering three different samples: Sample 1 (missing MRLs not replaced); Sample 2 (missing MRLs only replaced by default MRLs) and Sample 3 - the “full sample” - (missing MRLs in Sample 2 replaced by sample maxima).

5 Data

We use data on MRL regulation covering the period between 2005 and 2014 in the following 53 importing and exporting countries: Argentina, Australia, Brazil, Canada, Chile, China, Colombia, Egypt, India, Israel, Japan, Korea, Mexico, Malaysia, Norway, New Zealand, Russia, Singapore, South Africa, Switzerland, Thailand, Turkey, Ukraine, USA, Vietnam and 28 EU Member States. The data on MRL regulation were acquired from Agrobase-Logigram, a private company that maintains Homologa, the Gobal Crop Protection Database, using information from pertinent national ministries and legal publications.

However, the richness of the data received from Homologa that covers 2638 products[7] could not be fully exploited because a large amount of crops are too specific compared to the Harmonized System (HS) 6-level data. To enable an empirical trade analysis of these MRLs, it becomes impossible to use all the Homologa data since that would introduce MRL variation within the HS code that cannot be matched by trade variables. We therefore selected products that matched perfectly. These 31 products are reported in Annex Table 1.

The analysis is conducted at the disaggregated HS-six-digit product level, focusing on trade in HS Chapters 7 and 8 over 2005-14. These HS Chapters correspond to the agriculture fruit and vegetables sectors where pesticide MRLs are relevant. Fruits and vegetables in particular are interesting sectors to analyze because these are rejected more often than other products like meat or dairy products. For instance, the EU reports 2621 rejections by a member state of the EU from 2008 to 2015 with an increasing trend (Fiankor et al. 2016).

Export data come from the UN Comtrade database in current USD. Data on (simple average) applied tariffs are sourced from the International Trade Center. The bilateral trade cost variables are taken from CEPII (Head et al. 2010) and data for PTA-membership come from De Sousa[8].

Summary statistics are provided in Table 2. The full sample has more than 800,000 observations but export values are only positive for less than 14% of these. The original dataset without any MRL replacements (Sample 1) has more than 480,000 observations; the sample size goes up to 660,000 with missing MRLs replaced with default values (Sample 2), and further to 819,900 with sample maxima used to replace missing MRLs in cases which did not even report default MRLs (Sample 3).

[7] Including subcategories of products at various levels of aggregation.
[8] http://jdesousa.univ.free.fr/data.htm.

Table 2: Summary statistics

	Sample 1			Sample 2			Sample 3		
Variable	Obs	Mean	Std. Dev.	Obs	Mean	Std. Dev.	Obs	Mean	Std. Dev.
R_{ijpt}	480564	.336	.285	660666	.345	.296	807710	.346	.272
S^M_{ijpt}	569172	.099	.163	704112	.137	.201	808271	.17	.219
S^X_{ijpt}	580154	.112	.179	731634	.137	.203	811850	.17	.219
Exports (USD '000s)	580154	405.3	5782.8	731634	373.8	5589.8	811850	359.2	5651.3
Distance (km)	580154	6166.2	5049.8	731634	6155.6	5039.6	811850	6220.8	5010.4
Contiguity	580154	.049	.217	731634	.049	.215	811850	.048	.214
Common language	580154	.073	.260	731634	.069	.253	811850	.0673499	.2506271
Common colony	580154	.032	.177	731634	.031	.175	811850	.031	.173
PTA membership	580154	.519	.5	731634	.52	.5	811850	.501	.5
Common religion	580154	.166	.372	731634	.158	.364	811850	.158	.364
Tariffs (simple avg. appd., %)	580154	1.066	.224	731634	1.062	.212	811850	1.065	.206
Share of zero exports	84%			85%			86%		
Correlation between S^M,S^X	-0.13			-0.04			-0.13		

Source of variables: Exports (UN Comtrade); R_{ijpt}, S^M_{ijpt}, S^X_{ijpt} (Homologa); Tariffs (ITC); Distance, Contiguity, Common language, Common colony (CEPII, Head et al. 2010); Common religion (Helpman et al. 2008); PTA membership (De Souza).
Sample 1 (S1) only includes observations where the importer and the exporter had an explicit MRL.
Sample 2 (S2) = S1 + use of default MRLs to replace missing MRLs.
Sample 3 (S3) = S2 + use of sample maxima to replace those missing MRLs which also lacked a default MRL.

Available evidence in this literature suggests that harmonization within the EU has generally tended towards the high range of initial standards. For example, Vogel (2009) points out that the role of the EU's richest and most powerful members, which have traditionally imposed the strictest standards, has been critical in setting the standards agenda within the EU; their political and economic importance has served to make EU standards progressively stricter.

The Communities (1998) Single Market Review also concludes that the harmonized standards in most reviewed industries have been set higher than initial levels in most member countries. The history of EU automobile emission, chemical, and packaging standards also demonstrates that these standards have frequently been harmonized at levels slightly lower than those preferred by the EU's most stringent states (Germany, Denmark, Netherlands), but higher than those favoured by less stringent members (Italy, UK, and Spain) (Vogel, 2009).

This upward harmonization of standards is also true of the EU's MRL harmonization as of September 2008. A descriptive analysis of the pesticide MRL data for EU Member States for the full sample (Sample 3) before and after 2008 in Table 3 reveals that the MRLs have been harmonized at lower levels post-2008 i.e. the regulation has been made more stringent.

Thus, any increase in exports to EU destinations over 2009-2014 for the full sample despite stricter importer standards must surely be on account of harmonization.[9]

Table 3: Sample averages

	Intra-EU partners		**Non-EU exp to EU imp**	
	2005-2008	**2009-2014**	**2005-2008**	**2009-2014**
Average export propensity	0.306	0.307	0.075	0.071
Average export value ($)	712'199	714'933	86'681	121'228
Average R_{ijpt} (Sample 1)	0.001	0.000	0.024	0.015
Average R_{ijpt} (Sample 2)	0.001	0.000	0.019	0.016
Average R_{ijpt} (Sample 3)	0.003	0.000	0.019	0.030
MRL_Imp (Sample 1)	0.41	0.46	0.81	1.06
MRL_Imp (Sample 2)	0.37	0.39	0.76	0.93
MRL_Imp (Sample 3)	1.17	0.39	2.13	0.93
MRL_Exp (Sample 1)	0.41	0.46	2.99	2.83
MRL_Exp (Sample 2)	0.37	0.39	1.82	2.25
MRL_Exp (Sample 3)	1.17	0.39	2.71	2.69

Note: Sample 1 only includes observations where the importer and the exporter had an explicit MRL. Sample 2 = Sample 1 + use of default MRLs to replace missing MRLs. Sample 3 ("Full sample") = Sample 2 + use of sample maxima to replace those missing MRLs which also lacked a default MRL.

Finally, the sample averages in Table 3 also suggest that the period after MRL harmonization was associated with greater export value for both intra-EU and non-EU partners exporting to

9 This is also why we focus on the sub-component, S^{M}_{ijpt}, of R_{ijpt} where the importer is stricter than the exporter, in sensitivity analysis (see Section 7 for details).

the Common Market as well as a higher export propensity within the EU. These descriptive statistics suggest that MRL harmonization within the EU may have been associated with a positive trade effect at both margins. We explore this more formally in the next section.

6 Empirical model

Our empirical analysis is conducted in the framework of the gravity model as laid down by Anderson (1979). Following Anderson and van Wincoop (2004), the value of exports from country i to country j of product p at time t can be written as follows:

$$X^p_{ijt} = \frac{E^p_{jt} Y^p_{it}}{Y^p_t} \left(\frac{\phi^p_{ijt}}{P^p_{it} \Pi^p_{jt}} \right)^{(1-\sigma^p)} \tag{8}$$

where X^p_{ijt} denotes the value of exports of product p from country i to j at time t, E^p_j is the expenditure in the destination country j of product p, Y^p_i denotes the total sales of exporter i towards all destinations, Y^p is the total world output of product p, ϕ_{ij} are the bilateral trade costs and σ^p is the elasticity of substitution across products. P^p_{it} and Π^p_{jt} are the Multilateral Resistance Terms (MRTs) i.e. the outward and inward relative resistance of a country's exports towards all destinations and from all origins.[10] Since these terms are difficult to construct directly as national price indices are needed, applications of the gravity model resort to using dummy variables to control for them instead. At the sectoral level, time-varying importer-product and exporter-product fixed effects control for the MRTs in a panel setting (Anderson and Yotov, 2012).

Bilateral trade costs in ϕ_{ijpt}[11] arise from different sources such as import tariffs, τ_{ijpt}; geographical distance between trading partners, $\ln(Dist_{ij})$; cultural distance proxied by dummy variables identifying whether the trading partners share a common border, $Contig_{ij}$, had a colonial relationship, $Colony_{ij}$, and share a common language, $ComLang_{ij}$.

These variables enter ϕ_{ijpt} as follows:

$$\phi^{1-\sigma}{}_{ijpt} = exp(\beta_1 ln(1 + \tau_{ijpt}) + \beta_2 ln(Dist_{ij}) + \beta_3 Contig_{ij} + \beta_4 ComLang_{ij}) \tag{9}$$

Substituting (9) into (8), adding an error term, and taking the log of the resulting multiplicative model, yields the following estimating equation:

$$ln(X_{ijpt}) = \beta_1 ln(1+\tau_{ijpt}) + \beta_2 ln(Dist_{ij}) + \beta_3 Contig_{ij} + \beta_4 ComLang_{ij} + \mu_{ipt} + \gamma_{jpt} + \epsilon_{ijpt} \tag{10}$$

10 The MRTs are derived theoretically in Anderson and Van Wincoop (2003).

11 The notation, regarding the subscripts, is slightly modified to accommodate the product dimension, p.

where μ_{ipt} and γ_{jpt} are the time-varying exporter-product and importer-product fixed effects that proxy the MRTs and ε_{ijpt} is the error term.

In the context of this study, dyadic differences in MRL regulation, R_{ijpt}, can not only add to bilateral trade costs but the information disclosed by more stringent regulation can also enhance demand in the importing country by altering consumer preferences (for instance see Xiong and Beghin, 2014). Meanwhile, differences in MRL standards across EU Member States, (σ^{EU}_{pt}), is only likely to be associated with a trade cost effect.

Equation (10) is thus augmented to include both $\ln(1+R_{ijpt})$ and $\ln(1+ \sigma^{EU}_{pt})$. To compare the average trade effects over 2005-2008 and 2009-2014 i.e. in the pre- and post-EU MRL harmonization periods, we also include interaction terms $H*\ln(1+R_{ijpt})$ and $H*\ln(1+ \sigma^{EU}_{pt})$ along with H, which is a binary dummy variable that takes the value of 0 over 2005-2008 and the value 1 thereafter.

Finally, recent advancements in the estimation of structural gravity advocate the use of three-way fixed effects to mitigate endogeneity-induced biases in estimation (for instance see Baier and Bergstrand, 2007; Baier et al. 2014; Piermartini and Yotov, 2016; Shingal et al. 2017). The bilateral trade cost variables in equation (10) are thus subsumed in bilateral pair-wise fixed effects (χ_{ij}).

The final estimating equation is as follows:

$$ln(X_{ijpt}) = \beta_1 ln(1+R_{ijpt}) + \beta_2 H * ln(1+R_{ijpt}) + \beta_3 ln(1+\sigma^{EU}_{pt}) + \beta_4 H * ln(1+\sigma^{EU}_{pt}) + \beta_5 H + \beta_6 ln(1+\tau_{ijpt}) + \mu_{ipt} + \gamma_{jpt} + \chi_{ij} + \epsilon_{ijpt} \tag{11}$$

We also allowed exports to respond to regulatory heterogeneity with a lag. These results were found to be qualitatively similar and are available upon request.

6.1 Estimation issues

Two stylized features of trade data that challenge the estimation of structural gravity models are sample selection and heteroskedasticity (Xiong and Chen, 2014). In the agricultural trade sample we focus on, X_{ijpt} was found to equal 0 in 86% of all observations (details in Section 5). Sample selection was therefore clearly a concern with our data.

The Poisson Pseudo-Maximum Likelihood (PPML; Silva and Tenreyro, 2006) estimator is now regarded as the gold standard (for instance see Piermartini and Yotov, 2016) in the estimation of structural gravity models characterized by sample selection and heteroskedasticity. We therefore account for zero trade flows in the data using the PPML, which also addresses problems associated with heteroskedastic errors by characterizing trade multiplicatively in levels as opposed to log-linearly. Thus, we have:

$$X_{ijpt} = exp[\beta_1 ln(1+R_{ijpt}) + \beta_2 H * ln(1+R_{ijpt}) + \beta_3 ln(1+\sigma^{EU}_{pt}) + \beta_4 H * ln(1+\sigma^{EU}_{pt}) + \beta_5 H + \beta_6 ln(1+\tau_{ijpt}) + \mu_{ipt} + \gamma_{jpt} + \chi_{ij}] + \epsilon_{ijpt} \tag{12}$$

We also split our country sample into two sub-samples: (1) the sample of intra-EU trading partners; and (2) the sample of non-EU exporters accessing the Common Market. If regulatory heterogeneity is associated with a dominant trade cost effect, then we hypothesize MRL harmonization to be associated with a positive trade effect for both sub-samples.

7 Results and analysis

In this section, we report estimation results on the effects of the EU's MRL harmonization for each of the two sub-samples mentioned above.

7.1 Intra-EU trading partners

Table 4, panel A reports the baseline results using the PPML on all three samples for intra-EU trading partners in the data. All estimations include bilateral pair-wise and time- varying importer-product and exporter-product fixed effects, with the product defined at the HS-4 product level to reduce the dimension of the econometric specification and to obviate concerns about fixed effects constructed at the HS-6 product level being collinear with tariffs and the dyadic heterogeneity index. Since the dyadic heterogeneity index varies by dyad-product-year, standard errors are also clustered at that level.

Table 4: Baseline PPML estimates

	A: Intra-EU exports			**B: Non-EU exports**		
	(S1)	**(S2)**	**(S3)**	**(S1)**	**(S2)**	**(S3)**
	$\mathbf{X_{ijpt}}$	$\mathbf{X_{ijpt}}$	$\mathbf{X_{ijpt}}$	$\mathbf{X_{ijpt}}$	$\mathbf{X_{ijpt}}$	$\mathbf{X_{ijpt}}$
$\ln(1+R_{ijpt})$	-109.4***	-107.8***	-83.90***	-10.18**	-6.339*	-29.40***
	(35.98)	(35.22)	(12.81)	(4.595)	(3.351)	(7.560)
$H*\ln(1+R_{ijpt})$				10.85	-0.580	27.44***
				(10.10)	(8.114)	(7.666)
$\ln(1+\sigma^{EU}_{pt})$	-1.120***	-1.219***	-0.749***	-1.110***	-1.089***	-0.606***
	(0.161)	(0.170)	(0.0911)	(0.324)	(0.309)	(0.171)
$\ln(1+\tau_{ijpt})$	16.30***	17.29***	16.12***	0.904	1.040	0.939
	(4.422)	(4.339)	(3.863)	(0.856)	(0.913)	(0.753)
N	107,098	150,628	151,913	43,288	56,245	75,786
Pseudo-r2	0.787	0.789	0.789	0.754	0.756	0.761
Method			PPML 3wfe			
Fixed effects			ipt, jpt, ij			

Note: S1 only includes observations where the importer and the exporter had an explicit MRL. S2 = S1 + use of default MRLs to replace missing MRLs. S3 = S2 plus use of sample maxima to replace those missing MRLs which also lacked a default MRL. Product dimension in the fixed effects is at the HS4 digit level. Robust standard errors, clustered by dyad-product-year, included in parentheses. Variables H and $H*\ln(1+\sigma^{EU}_{pt})$ were dropped to achieve convergence. Levels of significance: *p<0.1 **p<0.05 ***p<0.01.

A priori, if the Common Market and the Cassis de Dijon principle were working perfectly, MRL harmonization within the EU should be redundant. Thus, we should not expect MRL harmonization within the EU to have any effect on intra-EU trade. In fact, even the coefficients of $\ln(1+R_{ijpt})$ and $\ln(1+\sigma^{EU}_{pt})$ in the results reported in Table 4, panel A, should be either economically or statistically indifferent from zero. However, the respective coefficients are found to be consistently negative across the three samples (see columns S1-S3 under "Intra-EU exports").

These results question the working of the Internal Market and the Cassis de Dijon principle within the EU. While $\ln(1+\sigma^{EU}_{pt})$ is zero by construction for intra-EU partners over 2009-2014, the negative coefficient of the variable in the pre-harmonization period across the three samples suggests that regulatory heterogeneity within the EU may have had an adverse effect on exports within the EU over 2005-2008 and that this effect could likely be mitigated, if not completely reversed, by the EU's MRL harmonization in September 2008.

7.2 Non-EU countries exporting to the EU

Panel B of Table 4 reports the baseline results on all three samples for non-EU countries exporting to the EU. Again, all estimations include time-varying importer-product and exporter-product fixed effects, with the product defined at the HS-4 product level to reduce the dimension of the econometric specification and to obviate concerns about fixed effects constructed at the HS-6 product level being collinear with tariffs and the dyadic heterogeneity index. Since the dyadic heterogeneity index varies by dyad-product-year, standard errors are also clustered at that level.

Relative MRL-stringency in a dyad has a strong negative effect on exports in the results reported in Table 4, panel B for all three samples; the coefficient of $\ln(1+R_{ijpt})$ is negative in all these cases. The coefficient of the interaction term $H*\ln(1+R_{ijpt})$ is found to be positive and statistically significant under column S3, "Non-EU exports". The magnitude of the coefficient suggests that the adverse effect of dyadic regulatory heterogeneity on non-EU exports to the EU over 2005-2008 may have been somewhat offset by the positive impact of the EU's MRL harmonization over 2009-2014.

Meanwhile, the coefficient of $\ln(1+\sigma^{EU}_{pt})$ is also found to be consistently negative across the three samples in the results reported in Table 4, panel B, pointing to the additional trade cost effect of different MRL regulation across EU Member States for non-EU exporters.

7.2.1 Decomposing non-EU exporters into OECD and non-OECD countries

Since our country sample of non-EU exporters comprises both OECD and non-OECD countries, we further examine the results reported in Table 4 by the level of economic development of non-EU countries exporting to the EU. To do so, we interact the variables of interest with a dummy variable OECD that takes the value one when a non-EU exporting country belongs to the group of OECD countries. These results are reported in Table 5, again for all the three samples.

Table 5: PPML estimates of non-EU exports (OECD vs non-OECD)

	Non-EU exports		
	(S1)	**(S2)**	**(S3)**
	X_{ijpt}	X_{ijpt}	X_{ijpt}
$\ln(1+R_{ijpt})$	-3.339*	-3.217	-24.87***
	(1.988)	(2.018)	(7.725)
$\ln(1+\sigma^{EU}_{pt})$	-1.068	-1.279**	-0.330*
	(0.724)	(0.605)	(0.188)
$H*\ln(1+R_{ijpt})$	-19.19*	-15.55*	22.60***
	(10.82)	(8.509)	(7.870)
$OECD*\ln(1+R_{ijpt})$	-29.75***	-26.54*	-8.880
	(11.52)	(15.90)	(18.41)
$H*OECD*\ln(1+R_{ijpt})$	54.54**	32.03	7.183
	(21.75)	(23.29)	(21.32)
$OECD*\ln(1+\sigma^{EU}_{pt})$	-0.351	-0.0316	-1.090***
	(0.780)	(0.672)	(0.290)
$\ln(1+\tau_{ijpt})$	-2.217***	-1.949***	-1.127*
	(0.737)	(0.721)	(0.627)
$OECD*\ln(1+\tau_{ijpt})$	13.56***	14.49***	14.10***
	(2.001)	(2.010)	(1.998)
N	43,288	56,245	75,786
Pseudo r2	0.760	0.761	0.766
Method		PPML 3wfe	
Fixed effects		ipt, jpt, ij	

Note: S1 only includes observations where the importer and the exporter had an explicit MRL. S2 = S1 + use of default MRLs to replace missing MRLs. S3 = S2 plus use of sample maxima to replace those missing MRLs which also lacked a default MRL. Product dimension in the fixed effects is at the HS4 digit level. Robust standard errors, clustered by dyad-product- year, included in parentheses. Variables H, $H*\ln(1+\sigma^{EU}_{pt})$ and $H*OECD*\ln(1+\sigma^{EU}_{pt})$ were dropped to achieve convergence. Levels of significance: *p<0.1 **p<0.05 ***p<0.01.

The coefficients of $\ln(1+R_{ijpt})$ and $\ln(1+\sigma^{EU}_{pt})$ are found to be negative in two of the three samples in these results, which points to the adverse effect of regulatory heterogeneity (both dyadic and intra-EU) on non-EU developing countries' exports to the Common Market in the pre-harmonization period. The coefficient of $H*\ln(1+R_{ijpt})$ is found to be positive under column S3; its magnitude suggests that the adverse effect of dyadic regulatory heterogeneity on non-EU developing country exports to the EU over 2005-2008 may have been somewhat offset by the positive impact of the EU's MRL harmonization over 2009-2014.

The coefficients of $OECD*\ln(1+R_{ijpt})$ and $OECD*\ln(1+\sigma^{EU}_{pt})$ are found to be negative and strongly significant under columns S1 and S3, respectively, which points to the adverse effect of regulatory heterogeneity (both dyadic and intra-EU) on non-EU OECD countries' exports to the Common Market in the pre-harmonization period. The coefficient of H*OECD*ln(1+ R_{ijpt}) is also found to be positive under S1; the magnitude of the coefficient suggests that the adverse effect of dyadic regulatory heterogeneity on non-EU OECD exports to the EU over 2005-2008 may have been more than offset by the positive impact of the EU's MRL harmonization over 2009-2014.

The results suggest that the overall findings in Table 4 are driven more by the sub-sample of non-EU OECD exporters.

7.2.2 Splitting regulatory heterogeneity into relative importer and exporter stringency

By construction, the index of regulatory heterogeneity, R_{ijpt}, includes both cases of relative importer and relative exporter stringency. However since sample averages reported in Table 3 reveal that MRLs within the EU have been harmonized to lower levels i.e. MRL regulation made more stringent, at least for the full sample, a positive coefficient of the interaction term H*ln(1+R_{ijpt}) in cases where the EU importer is stricter than the non-EU exporter would provide more conclusive evidence of the positive impact of MRL harmonization within the EU.

To enable this analysis, we split the index of regulatory heterogeneity, R_{ijpt}, (and the associated interaction terms) into two sub-indices S^M_{ijpt} and S^X_{ijpt}, denoting relative importer and exporter stringency respectively, as follows:[12]

At the pesticide level:

$$s^m_{ijpkt} = \begin{cases} \frac{abs(MRL_{ipkt}-MRL_{jpkt})}{MRL_{ipkt}+MRL_{jpkt}} & \text{if } MRL_{ipkt} > MRL_{jpkt} \\ 0 & \text{otherwise} \end{cases} \tag{13}$$

$$s^x_{ijpkt} = \begin{cases} \frac{abs(MRL_{ipkt}-MRL_{jpkt})}{MRL_{ipkt}+MRL_{jpkt}} & \text{if } MRL_{ipkt} \leq MRL_{jpkt} \\ 0 & \text{otherwise} \end{cases} \tag{14}$$

and at the product level:

$$S^M_{ijpt} = \frac{1}{K}\sum_{k=1}^{K} s^m_{ijpkt} \tag{15}$$

$$S^X_{ijpt} = \frac{1}{K}\sum_{k=1}^{K} s^x_{ijpkt} \tag{16}$$

The results from replacing R_{ijpt} by these sub-indices, reported in Table 6, confirm that the findings of Table 4, panel B are driven by relative importer stringency, especially under

12 In another robustness check, we experimented with a stronger definition of relative exporter stringency, i.e. $MRL_{ipkt} < MRL_{jpkt}$. Our empirical findings were found to be robust to this change in definition.

column (S3), thereby providing more conclusive evidence of the positive effects of the EU's MRL harmonization on market access to its non-EU partners. If non-EU countries export more to the EU despite MRL regulation in the EU becoming more stringent, then this must be on account of MRL harmonization.

Table 6: PPML estimates with dichotomized Rijpt

		Non-EU exports	
	(S1)	**(S2)**	**(S3)**
	$\mathbf{X_{ijpt}}$	$\mathbf{X_{ijpt}}$	$\mathbf{X_{ijpt}}$
$\ln(1+S^{M}_{ijpt})$	-10.79*	-6.782	-38.02***
	(6.329)	(4.180)	(9.387)
$H*\ln(1+S^{M}_{ijpt})$	-8.267	-40.75***	36.25***
	(13.16)	(13.99)	(9.474)
$\ln(1+S^{X}_{ijpt})$	-122.4***	33.29*	14.91
	(45.62)	(17.15)	(11.49)
$H*\ln(1+S^{X}_{ijpt})$	88.63*	-14.96	-9.993
	(50.61)	(24.93)	(18.86)
$\ln(1+\sigma^{EU}_{pt})$	-0.296	-0.210	-0.624***
	(0.232)	(0.245)	(0.166)
$\ln(1+\tau_{ijpt})$	0.202	0.890	1.021
	(0.698)	(0.758)	(0.741)
N	64,950	72,131	75,760
Pseudo r2	0.760	0.762	0.762
Method		PPML 3wfe	
Fixed effects		ipt, jpt, ij	

Note: S1 only includes observations where the importer and the exporter had an explicit MRL. S2 = S1 + use of default MRLs to replace missing MRLs. S3 = S2 plus use of sample maxima to replace those missing MRLs which also lacked a default MRL. Product dimension in the fixed effects is at the HS4 digit level. Robust standard errors, clustered by dyad- product-year, included in parentheses. Variables H and $H*\ln(1+\sigma^{EU}_{pt})$ were dropped to achieve convergence. Levels of significance: *p<0.1 **p<0.05 ***p<0.01.

When we replicate this analysis to examine the results for OECD and non-OECD countries amongst non-EU exporters, we still find the overall findings (see Table 7) to be driven by the sub-sample of non-EU OECD exporters *a la* Table 5.

Table 7: PPML estimates of non-EU exports (OECD vs non-OECD) with dichotomized Rijpt

	(S1) X_{ijpt}	(S2) X_{ijpt}	(S3) X_{ijpt}
$\ln(1+S^{M}_{ijpt})$	-3.658*	-3.097	-24.57***
	(2.180)	(1.913)	(8.389)
$\ln(1+S^{X}_{ijpt})$	-104.0**	-41.43*	-48.20
	(46.29)	(21.20)	(35.45)
$H*\ln(1+S^{M}_{ijpt})$	-43.56***	-52.04***	22.98***
	(15.89)	(17.30)	(8.490)
$H*\ln(1+S^{X}_{ijpt})$	25.95	26.56	28.25
	(51.42)	(28.17)	(38.72)
$OECD*\ln(1+S^{M}_{ijpt})$	-79.47***	-112.0***	-95.51***
	(21.02)	(24.50)	(20.93)
$OECD*\ln(1+S^{X}_{ijpt})$	62.81	130.5***	80.90**
	(156.0)	(41.19)	(40.89)
$H*OECD*\ln(1+S^{M}_{ijpt})$	109.8***	93.75**	66.21**
	(31.48)	(37.97)	(25.77)
$H*OECD*\ln(1+S^{X}_{ijpt})$	116.2	26.80	37.33
	(174.8)	(76.92)	(72.91)
$\ln(1+\sigma^{EU}_{pt})$	0.134	0.352	-0.339*
	(0.319)	(0.321)	(0.189)
$OECD*\ln(1+\sigma^{EU}_{pt})$	-1.311***	-1.616***	-1.022***
	(0.416)	(0.436)	(0.277)
$\ln(1+\tau_{ijpt})$	-2.087***	-1.212*	-1.230*
	(0.627)	(0.638)	(0.637)
$OECD*\ln(1+\tau_{ijpt})$	13.25***	13.45***	13.88***
	(1.947)	(1.965)	(1.992)
N	64,862	71,871	75,596
Pseudo-r2	0.766	0.768	0.767
Method		PPML 3wfe	
Fixed effects		ipt, jpt, ij	

Note: S1 only includes observations where the importer and the exporter had an explicit MRL. S2 = S1 + use of default MRLs to replace missing MRLs. S3 = S2 plus use of sample maxima to replace those missing MRLs which also lacked a default MRL. Product dimension in the fixed effects is at the HS4 digit level. Robust standard errors, clustered by dyad-product-year included in parentheses. Variables H, $H*\ln(1+\sigma^{EU}_{pt})$ and $H*OECD*\ln(1+\sigma^{EU}_{pt})$ were dropped to achieve convergence. Levels of significance: *p<0.1 **p<0.05 ***p<0.01.

Relative importer stringency is found to have an adverse effect on both OECD (see columns S1-S3) and developing country (see columns S1 and S3) non-EU exports to the EU in the pre-harmonization period in these results. In contrast, relative exporter stringency is found to have an adverse effect on developing country (see columns S1 and S2) non-EU exports but a positive effect on non-EU OECD (see columns S2 and S3) exports to the Common Market over 2005-2008. This is a significant finding, especially given the positive correlation between relative exporter stringency and institutional quality explored in Shingal et al. (2017).

These results further suggest that harmonization may have partially offset the adverse effect of relative importer stringency for both non-EU OECD (see column S1) and developing country exports to the EU (see column S3). Moreover, harmonization may not have altered the positive effect of relative exporter stringency for non-EU OECD exporters to the Common Market.

7.2.3 Decomposing estimated effects along margins of trade

The PPML estimator does not enable a decomposition of the estimated trade effect along the extensive and intensive margins. We therefore use the two-stage Heckman (1979) to disentangle the trade effects. In doing so, we acknowledge the exclusion restriction issue in Heckman-type estimations emphasized in the heterogeneous firm trade literature (for instance see Head and Mayer, 2013) and thus, closely follow Helpman et al. (2008) in our estimation strategy. This leads to the following "selection" and "outcome" equations (17 and 18, respectively):

$$Pr(X_{ijpt} > 0) = \beta_1 ln(1 + R_{ijpt}) + \beta_2 H * ln(1 + R_{ijpt}) + \beta_3 ln(1 + \sigma_{pt}^{EU}) + \beta_4 H * ln(1 + \sigma_{pt}^{EU}) + \beta_5 H + \beta_6 ln(1 + \tau_{ijpt}) + \delta_1 ln(Dist_{ij}) + \delta_2 Contig_{ij} + \delta_3 ComLang_{ij} + \xi EV_{ijpt} + \mu_{ipt} + \gamma_{jpt} + \epsilon_{ijpt} \tag{17}$$

$$ln(X_{ijpt} | X_{ijpt} > 0) = \beta_1 ln(1+R_{ijpt}) + \beta_2 H * ln(1+R_{ijpt}) + \beta_3 ln(1+\sigma_{pt}^{EU}) + \beta_4 H * ln(1+\sigma_{pt}^{EU}) + \beta_5 H + \beta_6 ln(1+\tau_{ijpt}) + \delta_1 ln(Dist_{ij}) + \delta_2 Contig_{ij} + \delta_3 ComLang_{ij} + \varrho \eta_{ijpt} + \psi z_{ijpt} + \mu_{ipt} + \gamma_{jpt} + \epsilon_{ijpt} \tag{18}$$

The Heckman two-step estimation involves running a Probit in stage one ("selection equation") on a dichotomous variable identifying non-zero exports between country i and j of product p at time t against the explanatory variables included in equation (17). Stage two of the Heckman ("outcome equation") is an OLS with the natural logarithm of exports as dependent variable on the same set of control variables as in stage one with the exclusion of at least one variable that should ideally affect trade only at the extensive margin.

Following Xiong & Beghin (2014), we used an indicator variable for common religion interacted with HS-4-digit fixed effects as the exclusion variable, EV_{ijpt}, in equation (17) to allow for heterogeneity across sectors in the self-selection process.[13] The predicted probabilities, $\rho_{\hat{i}jpt}$, from equation (17) were used to construct the inverse mills ratio[14], η_{ijpt}, which was

[13] We got weaker but consistent results by interacting the common religion variable with fixed effects computed at the HS-6 digit level. These results are available upon request.

[14] $\eta(\hat{\rho}) = \varphi(\hat{\rho}) / \Phi(\hat{\rho})$, where $\varphi(\cdot)$ and $\Phi(\cdot)$ are the standard normal density function and the standard normal cumulative function, respectively and $\rho_{\hat{i}jpt}$ are the predicted probabilities from the selection equation.

included in equation (18) to control for the selection bias. Following (Helpman et al., 2008), we also controlled for biases emanating from firm heterogeneity in equation (18) by including a cube polynomial of z_{ijpt} where $z_{ijpt} = \eta_{ijpt} + \rho_{\hat{i}jpt}$.[15]

The use of fixed effects in Probit estimations has been questioned due to the incidental parameters problem. Incidental parameters are nuisance parameters whose number increase as the sample size increases (Lancaster, 2000) and which bias estimates of coefficients derived from non-linear estimations such as the Probit. Using a Linear Probability model (LPM), estimated with OLS instead of a Probit, is equivalent and yields estimated probabilities within the unit interval (Wooldridge, 2010). We do not have a perfectly saturated model, but as Wooldridge (2010) reiterates, because we care about the partial effect of the explanatory variables on the response probability on average across the explanatory variables, then even if some estimated probabilities lie outside the unit interval it is not so important. We therefore resorted to using the LPM for estimating equation (17).

Table 8 reports the results of these two-stage estimations on all three samples for non-EU countries exporting to the EU. Again, all estimations include time-varying importer-product and exporter-product fixed effects, with the product defined at the HS-4 product level to reduce the dimension of the econometric specification and to obviate concerns about fixed effects constructed at the HS-6 product level being collinear with tariffs and the dyadic heterogeneity index. Since the dyadic heterogeneity index varies by dyad-product-year, standard errors are also clustered at that level.

[15] Following (Helpman et al., 2008), we do not use the normality assumption to recover η_{ijpt} and z_{ijpt} from the selection equation and instead work directly with the predicted probabilities, $\rho_{\hat{i}jpt}$.

Table 8: Adverse effects of regulatory heterogeneity within the EU on non-EU imports (Heckman estimates)

	(S1)		(S2)		(S3)	
	(1)	(2)	(3)	(4)	(5)	(6)
	$Pr(X_{ijpt}>0)$	$\ln(X_{ijpt})$	$Pr(X_{ijpt}>0)$	$\ln(X_{ijpt})$	$Pr(X_{ijpt}>0)$	$\ln(X_{ijpt})$
$\ln(1+R_{ijpt})$	-0.179***	-8.652***	-0.126***	-11.33***	-0.221***	-12.89***
	(0.0438)	(3.281)	(0.0353)	(3.525)	(0.0373)	(2.440)
$H*\ln(1+R_{ijpt})$	-0.123	-2.920	0.111***	5.934	0.191***	10.43***
	(0.0803)	(8.024)	(0.0413)	(7.235)	(0.0403)	(2.691)
$\ln(1+\sigma^{EU}_{pt})$	-0.0204***	-0.717***	-0.0181***	-0.673***	-0.0144***	-0.242*
	(0.00654)	(0.244)	(0.00505)	(0.240)	(0.00306)	(0.135)
$\ln(1+\tau_{ijpt})$	-0.0598***	-4.846***	-0.0380***	-4.461***	-0.0365***	-4.370***
	(0.0142)	(0.630)	(0.0124)	(0.584)	(0.0117)	(0.519)
$\ln(Dist_{ij})$	-0.0750***	-1.340***	-0.0787***	-1.451***	-0.0756***	-1.472***
	(0.00392)	(0.238)	(0.00323)	(0.235)	(0.00270)	(0.192)
$Contig_{ij}$	0.321***	2.162***	0.289***	2.130***	0.273***	1.221**
	(0.00850)	(0.821)	(0.00758)	(0.755)	(0.00654)	(0.612)
$Colony_{ij}$	-0.0882***	-0.162	-0.0832***	-0.0686	-0.0698***	0.299
	(0.00643)	(0.319)	(0.00551)	(0.297)	(0.00508)	(0.220)
$ComLang_{ij}$	0.0951***	-0.700**	0.0956***	-0.712**	0.0894***	-0.845***
	(0.00474)	(0.301)	(0.00421)	(0.297)	(0.00392)	(0.245)
EV_{ijpt}	0.00379***		0.00310***		0.00308***	
	(0.000453)		(0.000390)		(0.000337)	
η_{ijpt}		-1,593*		-1,690*		-4,910***
		(910.0)		(1,016)		(1,031)
z_{ijpt}		-22,617*		-23,768		-71,349***
		(13,576)		(15,444)		(15,504)
z^2_{ijpt}		21,762		22,744		69,697***
		(13,422)		(15,427)		(15,416)
z^3_{ijpt}		-7,815		-8,117		-25,466***
		(4,959)		(5,762)		(5,735)
N	122,855	9,630	167,749	11,898	206,542	14,976
r2	0.346	0.673	0.327	0.682	0.322	0.661
Method	LPM	OLS	LPM	OLS	LPM	OLS
Fixed effects	ipt, jpt	ipt, jpt	ipt, jpt	ipt, jpt	ipt, jpt	ipt, jpt

Note: Sample 1 only includes observations where the importer and the exporter had an explicit MRL. Sample 2 = Sample 1 + use of default MRLs to replace missing MRLs. Sample 3 = Sam ple 2 + use of sample maxima to replace those missing MRLs which also lacked a default MRL. The exclusion variable used in the selection equations is a dummy variable for common religion interacted with HS4 product fixed effects. Product dimension in the fixed effects is at the HS4 digit level. LPM = Linear Probability Model. Robust standard errors, clustered by dyad-product- year, included in parentheses. Levels of significance: *p<0.1 **p<0.05 ***p<0.01.

Relative MRL-stringency in a dyad has a strong negative effect on both the probablity of exporting and export value in the results reported in Table 8 for all three samples; the coefficient of $\ln(1+R_{ijpt})$ is negative in all these cases. In contrast, the coefficient of the interaction term $H*\ln(1+R_{ijpt})$ is positive and statistically significant at the extensive margin for S2 and S3, and at the intensive margin for S3 in columns (3), (5) and (6) of Table 8. The magnitudes of the interaction term coefficients suggest that the adverse effects on the probability of non-EU exports to the EU over 2005-2008 may have been somewhat offset by the positive impact of the EU's MRL harmonization over 2009-2014. These results on the positive effects of standards harmonization on trade at the extensive margin are consistent with the findings in Shepherd (2007).

Finally, the coefficient of $\ln(1 + \sigma^{EU}{}_{pt})$ is also found to be consistently negative across the three samples at both the extensive and intensive margins in the results reported in Table 8, pointing to the additional trade cost effect of different MRL regulation across EU Member States for non-EU exporters.

8 Conclusion

In September 2008, MRL regulation at the EU Member State level was replaced by Community- wide regulation, providing a near-natural experiment setting for analysis. In an original em- pirical contribution, this paper studies the effect of this MRL harmonization within the EU on its intra- and extra-EU imports using pesticides-MRL data for 53 exporting and import- ing countries over 2005-2014. We also embed regulatory heterogeneity in the Melitz (2003) framework in a theoretical contribution.

Regulatory heterogeneity in the pre-MRL-harmonized period of our analysis (2005-2008) is found to be associated with adverse effects on intra-EU exports, which questions the implementation of the Cassis de Dijon principle. Our findings also suggest that MRL harmonization may have made it easier for the EU's non-EU - both OECD and developing country partners to export to the Common Market though the effect is found to be stronger for OECD exporters.

Finally, since the effects of regulatory heterogeneity are expected at the extensive margin of trade, the effects of MRL harmonization are also primarily observed on the probability of trade, which is consistent with the findings in this literature (for instance see Shepherd, 2007).

Note that the EU's MRL harmonization may have also affected domestic producers as the MRLs became more stringent in some EU countries post-harmonization. Some domestic producers may have even been excluded from the market, with likely spill-over effects on trade. It would be useful to examine these effects in future research on this subject.

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Annex

Annex Table 1: List of included products

HS Code	Product	HS Code	Product	HS Code	Product	HS Code	Product
080211/2	Almonds	080920	Cherries	080710	Melons	080430	Pineapples
080810	Apples	080240	Chestnuts	100820	Millet	080940	Plums
080910	Apricots	070320	Garlic	071120	Olives	081020	Raspberries
070920	Asparagus	080221/2	Hazelnuts	070310	Onions	070970	Spinach
070930	Aubergine	081050	Kiwi	080510	Oranges	081010	Strawberries
080440	Avocados	080530	Lemons	080720	Papayas	080231/2	Walnuts
070410	Broccoli	080520	Mandarins	080930	Peaches	080711	Watermelons
070940	Celery	080450	Mangos	080820	Pears		

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