Received: 29 September 2017 | Revised: 13 November 2018 | Accepted: 15 November 2018

DOI: 10.1111/twec.12756

ORIGINAL ARTICLE

WILEY The World Economy

Trade barriers or trade facilitators? The heterogeneous impact of food standards in international trade

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Funding information
Research was funded by the Research Council of Norway, project 216742 "Non-tariff barriers, food safety and international food trade" and project 233836 "Traders in the Food Value Chain: Firm Size and International Food Distribution".

KEYWORDS
food standards, international trade, Norway, seafood, WTO SPS agreement

1 | INTRODUCTION

International food trade is characterised by the growing importance of non-tariff measures (NTMs) such as sanitary and phytosanitary (SPS) measures and technical barriers to trade (TBT) related to health or food quality (WTO, 2012). Using recent, precise data collected for two dozen countries, Gourdon and Nicita (2012) found that 60% of trade in food-related products was affected by SPS measures. According to the WTO SPS and TBT agreements, such measures should not unnecessarily restrict trade. Despite this, concerns have been raised that NTMs are sometimes used as a trade policy measure (Gourdon & Nicita, 2012; WTO, 2012), and at times, foreign standards can have the same effect as severe trade costs. With some inspiration from early literature such as Deardorff and Stern (1998) and international data collection exercises, efforts were made to estimate the tariff equivalents of NTMs, see example, Kee, Nicita, and Olarreaga (2009). This contributed to consolidating a "standards-as-barriers" approach, where NTMs are technically considered as resembling tariffs.

Contrary to this, some authors have argued that foreign standards may sometimes promote export. There may be asymmetric information, whereby consumers have less information about product quality and safety than do the producers, and standards may reveal this information thereby increasing demand. For example, Chen, Wilson, and Otsuki (2008) investigated the impact in 17 developing countries of standards imposed by developed countries and found that quality standards tended to increase export, whereas certification procedures had more negative effects.

Effects can also differ across subsectors. Disdier, Fontagné, and Mimouni (2008) found that even if SPSs on the whole can be said to have a trade-reducing effect, there was a trade-enhancing effect for eight out of 30 subsectors of agricultural trade. Consequently, in the recent literature on NTMs and food standards, the "standards-as-barriers" perception is gradually challenged by the "dual face of standards" approach; even if there is a cost involved in complying with standards, their trade-enhancing effects may at times be even larger.

The first aim of this paper was to present a theoretical foundation for the "dual face of standards" hypothesis. I present a slight modification of Chaney's (2008) augmentation of the seminal Melitz (2003) trade model with heterogeneous firms. In contrast to those models, here trade costs are endogenous and depend on standard-compliance costs. In addition, there is a parameter in the utility function comprising a potentially positive link to standard compliance. This modelling set-up enables me to define two effects from foreign standards and study how they interact: a cost-increasing effect and a demand effect. Standards will be trade-enhancing if the latter effect more than cancels out the first one. This will happen if one or more of the following conditions are sufficiently strong: demand is sensitive to standard compliance, compliance costs are small, and compliance costs constitute a small part of the total. A few other theoretical models comprise these two effects, but in different competitive environments than here. In Ganslandt and Markusen (2001), there is oligopolistic competition, and in Chen et al. (2008), there is perfect competition. Here, in contrast, I work with monopolistic competition, which fits well with the empirical study to be undertaken (see footnote 23 below). Only Chen et al. (2008) comprise multiple countries as here.[1]

Some evidence indicates that NTMs are particularly prevalent for seafood. For example, Jaud, Cadot, and Suwa-Eisenmann (2012) found that for SPS sanitary risk alerts in the EU during 2001–05, 23% concerned unprocessed fish and seafood alone. Disdier and van Tongeren (2010) found that NTMs covered a large share of seafood trade, but there were few trade frictions or concerns. Shepotylo (2016) explored SPS and TBT notifications on seafood among WTO members and found that SPSs affected 15.4% of all possible bilateral seafood exports, while TBTs affected 3.1%. This high incidence of NTMs for seafood and the fact that seafood products are an important subset of international food trade, representing about 10% of global food trade, motivates further study of international trade in seafood.[2]

The second aim of this paper was therefore to study the impact of food standards for seafood export. The country studied is Norway—a big actor in the international seafood market: 90%–95% of its seafood production is exported,[3] and it is currently the world's second-largest exporter of seafood, accounting for 9% of world exports in 2010–12. I investigate how Norwegian exports of different seafood products to different countries are affected by foreign NTM measures, using data on WTO notifications of SPSs. Some earlier survey evidence revealed that adaptation to foreign standards was a very important part of export costs among Norwegian seafood exporters (Medin & Melchior, 2002). For about 20 countries, with the United States, Russia and Brazil on top, the exporting firms had experienced that veterinary standards had hindered exports (Medin & Melchior, 2002). More recently for Russia, Norwegian salmon exports and later other seafood products were subject to a more restrictive SPS regime (Holm &

[1]Also, Reyes (2011) discusses standards in a model based on Chaney (2008; a three-country version). However, he does not discuss the effect of a unilateral imposition of a standard, like here, but rather the effect on a third country of two countries harmonising their standards. In his model, standards act as fixed trade costs, whereas here, they can represent both fixed and variable costs.

[2]Using data from WITS/COMTRADE for 2010, global food imports (Harmonised System, HS, chapters 1–24) were US$1,141 billion, whereas seafood imports were US$110 billion, equivalent to 9.6%.

[3]This information was provided orally by the Norwegian Seafood Council.

Kokkvold, 2007).[4] Another case is China, where extended veterinary inspections were introduced after the controversial Norwegian award of the Nobel Peace Prize to a Chinese dissident in 2010 (Chen & Garcia, 2015).[5,6]

In line with some recent articles (Chen et al., 2008; Fontagné, Orefice, Piermartini, & Rocha, 2015), the basis for my analysis is firm-level data, which enables distinguishing between the *extensive* margin of trade (the number of firms exporting a particular product to a particular country) and the *intensive* margin (their average exports). I find that on average for all products, SPSs have a significant trade-reducing effect for both margins of trade.[7] Despite the negative overall effects, there are strong differences across subsectors: For fresh seafood products, the negative effect is cancelled out by a counteracting positive effect. This study therefore adds to the literature suggesting that the impact of NTMs on trade varies across products, and NTMs in general cannot be seen as trade-reducing tariff equivalents.

Section 2 presents the theoretical background and Section 3 the empirical evidence. Section 4 concludes.

2 | FOOD STANDARDS, FIRM HETEROGENEITY AND THE EXTENSIVE AND INTENSIVE MARGINS OF TRADE

This section provides a theoretical foundation for the "dual face of standards" hypothesis. The model presented is a slightly modified version of Chaney's (2008) augmentation of the Melitz (2003) trade model, and the reader is referred to those articles for full details on the model. The model differs from Melitz/Chaney in that export costs contain elements reflecting costs of compliance with food standards imposed by an importer. Furthermore, demand in an importing country may increase as a consequence of such compliance. I examine under which circumstances a standard will be export-reducing and under which it will be export-enhancing, and I separate the effects on total export into extensive and intensive margins.

2.1 | The model

There are M countries in the world exporting each other. In the empirical analysis in Section 3, we study one single country that exports to many destinations, and thus, the model will mainly be presented from this viewpoint. Utility in a given country is a classical two-level function, where the upper level is a Cobb–Douglas aggregate of a homogenous good and various differentiated product groups (often also referred to as sectors). The focus is on any of the differentiated product groups,

[4]Furthermore, from August 2014, all Norwegian seafood exports to Russia were banned due to the sanctions related to the Ukraine crisis.

[5]The case was also covered in reputable media. See, for example, the Norwegian web-based newspaper E24. 19 September 2012: "Norsk laks kan bli liggende 14 dager på flyplassen." https://e24.no/makro-og-politikk/kinas-nobel-raseri-rammer-norsk-fisk-knallhardt/20276015.

[6]Note, however, that these cases are not represented in the data used for regression analysis. Russia because it was not a WTO member until 2012 and is therefore not included in the SPS data set, which only includes years up to 2011; and China because the notification data do not count unofficial measures like those imposed on Norway after 2010.

[7]Using data for aggregated trade flows, Shepotylo (2016) found that foreign SPSs were often positively associated with the probability of export of a given seafood product to a given country while negatively associated with the traded value. However, there was considerable heterogeneity in responses across products. Fontagné et al. (2015) studied how export decisions of French firms were affected by unresolved SPS concerns and found that they had a negative impact on the probability of export as well as on the export value.

and the lower-level sub-utility function for a group like that, U, is given by a CES aggregate of the different varieties available within the group:

$$U = \left(\int_{\omega \in \Omega} (qx(\omega))^{\frac{\sigma-1}{\sigma}} \mathrm{d}\omega \right)^{\beta \frac{\sigma}{\sigma-1}}, \tag{1}$$

β denotes the expenditure share for the product group, $\sigma > 1$ the elasticity of substitution between varieties within the group, $x(\omega)$ consumption of variety ω and Ω the set of available varieties. q is a quality-perception parameter related to food standards, and $q > 1$ implies that consumers get higher utility from consuming a product group if it complies with a food standard imposed.[8,9] Note that the size of the different parameters can differ across product groups, but since sub-utility of any of the groups is given by Equation (1) and the expenditure shares are constant due to the upper-level Cobb–Douglas function, it is sufficient to analyse one group. From now on, a group like that will be referred to as a *product*, while varieties within the group will be referred to as a *variety*.

Each firm produces one single variety under constant returns to scale. The only input in the economy is labour. Marginal production costs, c, differ across firms and are randomly drawn from a probability distribution with density given by $g(c)$.[10] Firms also face two types of export costs (equal for all firms within the same country): a fixed cost, $f > 0$; and a variable cost of the iceberg type, $t \geq 1$, where unity indicates no costs.[11] They depend on the country of production (j) as well as that of consumption (i). The costs faced by the exporter of complying with a standard imposed by the importing country can form part of both these costs. For example, if the importing country sets a stricter minimum acceptable bacteria level in salmon, the exporter may have to improve cooling facilities during storage as well as transportation. As a consequence, the exporter may have to invest in a new storage (increased fixed compliance costs) or pay more for transportation (increased variable compliance costs). F_{ji} and T_{ji} are, respectively, the part of f_{ij} and t_{ij} that are due to compliance costs, whereas φ_{ji} and τ_{ji} capture all other export costs (fixed and variable, respectively). f_{ij} and t_{ij} are given by:

$$f_{ji} = F_{ji}^{\lambda_{ji}} \varphi_{ji}^{1-\lambda_{ji}}, \tag{2}$$

$$t_{ji} = T_{ji}^{\mu_{ji}} \tau_{ji}^{1-\mu_{ji}}, \tag{3}$$

F_{ji} and T_{ji} are equal to or larger than 1, where unity indicates no compliance costs. The parameters λ_{ji} and μ_{ji} indicate the amount of the total trade cost (fixed and variable, respectively) that is due to compliance costs, and they lie between zero and one. q in Equation (1) is specific both to the exporting and the importing country, as it is linked to compliance costs by the following formula:[12]

[8]See, for example, Venables (1987) for a similar formulation of utility.

[9]Note that I assume that standards apply for the whole product group, not single varieties. This formulation fits the data used in the empirical part, where most seafood standards are given at the four-digit HS level (see Section 3.1).

[10]Varieties produced by firms with the same marginal cost are symmetric, thus it is sufficient to index firms by c, and we disregard ω in the following.

[11]Note that these costs only apply for $j \neq i$, which is the situation of interest here (all sales of interest are exports, and all standards of interest are imposed by a foreign country). Note, however, that the model also applies for $j = i$, and it would be straightforward allowing f and t to be larger than, respectively, zero or one also for $j = i$.

[12]The modelling of q is partly inspired by articles on firm-level heterogeneity in quality, such as Baldwin and Harrigan (2011).

$$q_{ji} = T_{ji}^{\varepsilon_{ji}} F_{ji}^{\eta_{ji}}, \tag{4}$$

$\varepsilon_{ji} \geq 0$ and $\eta_{ji} \geq 0$. Complying with a food standard increases demand in the importing country if ε_{ji} and η_{ji} are strictly positive, and large values indicate high elasticity of demand with respect to standard compliance. T_{ji} and F_{ji} can be viewed as costs of actual quality upgrades, for example, better cooling arrangements, or they can be viewed as investments in schemes that reduce uncertainty or improve the reputation and attractiveness of products; a better reputation may reduce the probability of being subjected to costly and tedious border controls where there is a food standard.

In accordance with Chaney (2008), there is a large mass of n_j producing firms in country j that is exogenous and proportional to the income/labour force. These engage in monopolistic competition, and there is free entry into exporting; thus, the consumer price of a given variety in country i will be a constant markup over marginal production costs, adjusted for t_{ji} and q_{ji}:[13]

$$p_{ji} = \frac{\sigma}{\sigma - 1} \frac{t_{ji} c}{q_{ji}}. \tag{5}$$

Maximisation of (1) yields demand faced by a firm equal to: $x_{ji} = t_{ji} p_{ji}^{-\sigma} B_i$. Using this together with Equations (5), (3) and (4), we can write an individual firm's revenue from exporting to i as:

$$r_{ji} = \left(T_{ji}^{\varepsilon_{ji} - \mu_{ji}} \tau_{ji}^{-(1-\mu_{ji})} F_{ji}^{\eta_{ji}} c \right)^{\sigma - 1} B_i \left(\frac{\sigma - 1}{\sigma} \right)^{(\sigma - 1)}, \tag{6}$$

where B_i reflects demand parameters:

$$B_i = \frac{\beta_i y_i}{P_i^{1-\sigma}}, \tag{7}$$

P_i is the product's contribution to the ideal price index. From Equation (1) this is given by:

$$P_i = \left[\sum_{j=1}^{M} n_j \int_0^{\bar{c}_{ji}} p_{ji}(c)^{1-\sigma} g_j(c) \mathrm{d}c \right]^{\frac{1}{1-\sigma}}. \tag{8}$$

Furthermore, and in accordance with Chaney (2008), countries are considered small enough not to influence y_i and P_i of their trading partners; thus, B_i is considered exogenous.

From Equation (6), we can find the elasticities of r_{ji} with respect to F_{ji} and T_{ji}:

$$E_{F_{ji}}(r_{ji}) = \eta_{ji}(\sigma - 1), \tag{9}$$

$$E_{T_{ji}}(r_{ji}) = (\varepsilon_{ji} - \mu_{ji})(\sigma - 1). \tag{10}$$

Firms with high marginal costs will not earn enough from exporting to country i to cover f_{ji}. Hence, they will not export there. A firm with a sufficiently low c to just find it profitable to export to i will, for its part, use all its export revenue from i to cover f_{ji}. The marginal cost of a firm like that, $\bar{c}_{ji}$, is called the cut-off cost, and it is found by setting $r_{ji}(\bar{c}_{ji}) = f_{ji}$ in Equation (6) and using Equation (2):[14]

[13]Equation (5) demonstrates the fact that in this type of models, q_{ji} works in the exact opposite direction of t_{ji} (see Venables, 1987).

[14]Firms with $c < \bar{c}_{ji}$ will earn positive profits from selling to country i.

$$\bar{c}_{ji} = T_{ji}^{\varepsilon_{ji}-\mu_{ji}} F_{ji}^{\frac{\eta_{ji}(\sigma-1)-\lambda_{ji}}{\sigma-1}} \tau_{ji}^{-(1-\mu_{ji})} \varphi_{ji}^{-\frac{1-\lambda_{ji}}{\sigma-1}} \frac{\sigma}{\sigma-1} B_i^{\frac{1}{\sigma-1}}. \tag{11}$$

From Equation (11), we can find the elasticities of $\bar{c}_{ji}$ with respect to F_{ji} and T_{ji}:[15]

$$E_{F_{ji}}(\bar{c}_{ji}) = \frac{\eta_{ji}(\sigma-1)-\lambda_{ji}}{\sigma-1}, \tag{12}$$

$$E_{T_{ji}}(\bar{c}_{ji}) = \varepsilon_{ji} - \mu_{ji}. \tag{13}$$

Aggregating over all firms, the extensive margin (the number of exporters) is given by:

$$n_{ji} = n_j \int_0^{\bar{c}_{ji}} g_j(c)\mathrm{d}c. \tag{14}$$

Total export is given by:

$$R_{ji} = n_j \int_0^{\bar{c}_{ji}} r_{ji}(c) g_j(c)\mathrm{d}c. \tag{15}$$

The intensive margin (average exports per firm) is given by:

$$I_{ji} = \frac{R_{ji}}{n_{ji}}. \tag{16}$$

The signs of Equations (9), (10), (12) and (13) are uncertain and depend upon the size of ε_{ji}, η_{ji}, μ_{ji} and λ_{ji}. To discuss the circumstances under which the expressions have a certain sign, I divide effects into cost and demand effects. The "cost effect" refers to the fact that a food standard requires firms to bear compliance costs. It is captured by the parameters reflecting the amount of total trade costs that is due to such costs (μ_j and λ_{ji}). This effect is also present in the standard Melitz–Chaney model.[16] The introduction of the parameter q_{ji} and its link to compliance costs, however, creates a demand effect not captured by that model. This "demand effect" refers to the fact that the imposition of a food standard increases demand in the importing country, and it is captured by the parameters reflecting how much demand increases due to compliance with the food standard (ε_{ji} and η_{ji}).

The demand effect dominates when ε_{ji} and/or η_{ji} are large or when μ_{ji} and/or λ_{ji} are small. The imposition of a food standard leads to an increase in the cut-off cost (Equations (12) and (13) are positive) and an increase in each firm's export (Equations (9) and (10) are positive). More firms find it profitable to export since the cut-off cost has increased. Thus, both total export and the extensive margin increase. The effect on the intensive margin is indeterminate.[17] In contrast, the cost effect dominates when μ_{ji} and/or λ_{ji} are large or when ε_{ji} and/or η_{ji} are zero or small. In this

[15]We implicitly assume that firms consider T_{ji} and F_{ji} as exogenous. Hence, firms do not choose values of these variables to upgrade the quality or reduce the uncertainty about their products.

[16]In fact, the model reduces to the Melitz–Chaney model if compliance costs (variable as well as fixed) are the only trade costs, and standards have no effect on demand (i.e., $\mu_{ji} = \lambda_{ji} = 1$ and $\varepsilon_{ji} = \eta_{ji} = 0$).

[17]The new exporters export less than the least-selling incumbent exporters (due to their higher marginal costs), and thus, the effect on average export per firm can go in either direction.

case, Equations (12) and (13) are negative, and the imposition of a food standard leads to a decrease in the cut-off cost. Total export and the extensive margin decrease, contrary to what was found for a dominating demand effect.[18] Consequently, any trade-reducing effects from standards may be offset if the imposition of a standard has a positive effect on demand. Therefore, the model provides a theoretical foundation for the "dual face of standards" hypothesis.

2.2 | Differences between countries and products

Whether cost or demand effects will dominate depends on characteristics of countries and products. The parameters T_{ji}, F_{ji}, ε_{ji}, η_{ji}, μ_{ji} and λ_{ji} may vary, for instance due to differences in preferences, technology, knowledge, reputation and product attributes. Norway is a highly developed country with high domestic standards and access to advanced technology, and thus, compliance capacity is likely to be high. If a new foreign standard is imposed, Norwegian products may already comply with it, or only minor increases in T_{ji} and/or F_{ji} may be required. This, together with good knowledge about foreign rules and laws among Norwegian exporters can make compliance costs constitute a small part of the total (low μ_{ji} and/or λ_{ji}). However, demand for the standard-imposed product may still increase, as the standard can give foreign buyers greater assurance about the product's quality and safety (ε_{ji} and/or η_{ji} are positive).

For fresh food, transport costs are particularly high, as this requires costly cooling arrangements during transport, and air shipment is likely. Then cost of compliance may constitute a small part of the total (low μ_{ji} and/or λ_{ji}). In addition, uncertainties about quality and safety are probably more pronounced for fresh food, and demand may be more sensitive to reputation and quality assurance (high ε_{ji} and/or η_{ji}). The Norwegian seafood industry has made large investments in generic marketing that extols the high quality, clean and healthy aspects of Norwegian seafood, and the products have a good reputation in many countries. We could therefore expect that Norway has a particularly high compliance capacity for seafood and especially for fresh products (low T_{ji} and/or F_{ji}). All these factors increase the likelihood of a positive demand effect for Norway, especially for fresh seafood, and thus, export may increase when a new foreign standard is imposed.[19] In the next section, I test this prediction empirically.

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[18]As long as compliance costs are mainly variable, each firm's export will also decline (Equation (10) is negative), and the effect on the intensive margin is indeterminate just as in the case of a dominating demand effect. If compliance costs are fixed, however, each firm's export will either not be affected (for $\eta_{ji} = 0$, F_{ji} is not part of Equation (6) and (9) is zero) or increase slightly (Equation (9) is slightly positive). This, together with the fact that the least-selling exporters stop exporting, leads to an increase in average export per firm, and thus, the effect on the intensive margin is positive. A more formal discussion of this phenomenon can be found in Lawless (2010).

[19]Whereas Norway and other developing countries are likely to have a high compliance capacity, the opposite may be true for many developing countries. This may distort demand away from those countries, further strengthening the positive effect on export for countries such as Norway. A distortion effect like this is present in the model: with a decrease in $\bar{c}_{ji}$ in a large number of other (developing) countries, P_i and hence B_i can no longer be considered exogenous. The number of firms from other countries exporting to i will decrease, and this will induce an increase in P_i and hence B_i (see Equation (7)). From Equations (6) and (11), the elasticities of $\bar{c}_{ji}$ and r_{ji} with respect to B_i are, respectively, $E_{B_i}(\bar{c}_{ji}) = \frac{1}{\sigma-1}$ and $E_{B_i}(r_{ji}) = 1$. These are positive. Consequently, even with no dominating demand effect in a country with high compliance capacity, the country may experience an increase in exports following the imposition of a new foreign standard. Such distortionary effects of standards were discussed by Disdier et al. (2008), who found that intra-OECD agricultural export was not impeded by foreign SPSs and TBTs, whereas developing country exports to the OECD were indeed affected. Similarly, Xiong and Beghin (2014) found that exporters from low-income countries were more constrained by a standard imposed by high-income OECD countries than were exporters from high-income countries.

3 | EMPIRICAL EVIDENCE: NORWEGIAN SEAFOOD EXPORTS AND THE ROLE OF STANDARDS

This section aims to provide new econometric evidence on the impact of SPSs in international seafood trade. Based on the theory framework presented in Section 2, a gravity analysis for Norwegian seafood export is performed, hypothesising that export is a function of the importers' GDP level and the costs of exporting.[20]

3.1 | Data and variables used in the regression analysis

I use a panel data set for Norwegian seafood exports at the firm level for years 2002–11, provided by Statistics Norway. Exporters include fish-farming producers, fish companies based on catch, seafood processing firms and pure trading companies. Export is disaggregated by destination country and products,[21] and the data reveal that the Norwegian seafood export business is relatively fragmented: there are above 400 exporters each year, and throughout the period, these sell 111 products to practically all countries in the world. In 2011, total Norwegian seafood exports stood at 6.8 billion euro, of which more than half (3.7 billion) were fresh products. Export of farmed salmon and trout products has been sky-rocketing in the later years and constituted 3.8 billion euro in 2011; the single largest product was fresh farmed salmon, which represented almost half of total seafood exports. In the recent years, there has been little export Russia due to the Ukrainian crisis, but in 2011 Russia was the most important export market for Norwegian seafood, closely followed by France. In addition to other EU countries (in particular Denmark and Poland), Japan and China were important.

For use as dependent variables, I construct three aggregated variables reflecting export to each country–product combination. These are total export value ($export_{ivt}$), the number of firms exporting ($extensive_{ivt}$) and their average export value ($intensive_{ivt}$), i denoting destination country and v product (not variety, as described in the theory part).[22,23] Export values are measured in current NOK. Following the tradition in the gravity literature, the logs of the variables are used in the regression.

[20]See UNCTAD and WTO (2012) for an introduction to gravity analysis.

[21]In the original data, products are given at the eight-digit HS level. For this analysis, I aggregate up the six-digit level to be able to match the export data with data from other sources. However, results from the regressions are robust to keeping the export data at the eight-digit level. Some products have also been further aggregated to get a consistent product classification during the period.

[22]I use aggregated variables as it reduces the influence of data noise, which is generally higher the more disaggregated the data are. For example, if a firm changes identification number due to data inaccuracy but the underlying activity is unchanged, this will appear as entry and exit in the microdata, but the aggregate variables (number of firms, average sales) will be unaffected. Furthermore, the proportion of zeros in the data is reduced. Also, other studies follow this approach (Buono & Lalanne, 2012). An alternative is to use the firm-level data directly, as in Chen et al. (2008) or Fontagné et al. (2015). This has the advantage of more accurately reflecting entry and exit, given that the data are correct. (By definition, firm-level data reflect gross exit and entry, whereas aggregate data on the number of firms reflect only net entry, unless special variables on entering and exiting firms are created.) Disaggregated data also yield a larger number of observations.

[23]In the theoretical model underlying the analyses presented here, it is assumed that firms are monopolistically competitive (see Section 2). This seems as a reasonable assumption. The Herfindahl–Hirschman index (HHI) for firm export value ranges from 140 to 431 in the different years. Calculated instead at the firm-product level, it ranges from near 0 to 216 for the different years and products (the possible range of HHI being from 0 to 10,000). In the sample used for regression, the number of firms exporting a given product ranges from 31 to 197 across different products and years.

For data on food standards, I use data on notifications of regular SPSs from the WTO Integrated Trade Intelligence Portal (I-TIP database).[24] It should be noted that these data do not measure the strictness or restrictive impact of SPSs, just the prevalence. The data also have other drawbacks: the number of notifications has increased considerably over time, which probably reflects notification practice rather than policy change. Moreover, withdrawals of SPSs are not reported. In addition, the quality of notifications in terms of accuracy and completeness has been questioned (Bacchetta, Richtering, & Santana, 2012). However, some of the potential "notification bias" originating from these issues is being limited by the relatively short time period of our analysis. There is also little reason to believe that heterogeneity across products—which is a main focus of this paper—should be affected by the bias. Since alternative sources of data on NTMs also have serious drawbacks, I choose to work with these data despite the issues discussed above.[25]

I construct a dummy variable for SPS notifications, sps_{ivt}, which is equal to 1 from year t and onwards if importing country i imposed at least one SPS on product v in year t. Most notifications are reported at the four-digit HS level for each country, but some are also reported at the two or six-digit levels.[26] When reported at a more aggregate level than six-digit level, I set the sps_{ivt} equal to 1 for all six-digit categories within the more aggregated category. Forty per cent of the observations included in the main analysis presented below are covered by at least one SPS. As argued in Section 2, the effect of standards may vary across products, and standards may have a greater impact on demand for fresh products. I therefore interact sps_{ivt} with a dummy for fresh seafood products (*fresh*). sps_{ivt} and its interaction with *fresh* serve as our main explanatory variables of interest.

I also include applied ad valorem tariffs against Norway for each importing country and product at the six-digit HS level, $tariff_{ivt}$. These data are taken from the MacMap database, which unfortunately only provides information for years 2001, 2004 and 2007.[27] To be able to take advantage of full time series of the export data, I linearly interpolate $tariff_{ivt}$ for the interim missing years and set the value for years after 2007 equal to the 2007 level. This should be a fairly good approximation, due to the stepwise reduction in tariffs agreed upon in the Uruguay Round and the fact that all commitments of tariff reductions were completed by 2005. Nevertheless, I perform three different sensitivity analyses of the treatment of the tariff variable, and they all show that the

[24]According to the WTO SPS agreement, governments implementing a new SPS measure must notify the WTO Secretariat, which makes the information publically available, that is, through the I-TIP database (http://i-tip.wto.org/goods/default.aspx?language=en). Also, see Shepotylo (2016) for a description of the data.

[25]An alternative would be to use data on concerns about SPS measures raised in the WTO SPS Committee, so-called STCs, as do Besedina (2015) and Fontagné et al. (2015; see the former for a detailed description of the data). However, as argued by Fontagné et al. (2015), the STC-SPSs are likely to be mainly trade restricting, as they relate to actual complains made by exporting countries towards importing countries. Therefore, the STC data are not well fit for analysis of the trade-enhancing vs trade-reducing effects of food standards, which is the focus in the present article.

[26]Sometimes, a product description is reported, but not the corresponding HS classification. In this case, I use the WTO interpreted HS category.

[27]The data also have another weakness: Tariffs for the new EU member states prior to their entry into the EU in 2004 and 2007 are erroneously set equal to the EU level. Except for Malta and Cyprus, Norway had free trade agreements with these countries prior to their joining the EU, so I change their tariffs to zero. I exclude all other countries for which tariffs are lacking some years. I choose to use this database despite its weaknesses because it also has strong advantages: It is constructed for analytical purposes and thereby contains better information on applied tariffs than do other sources. It also contains the ad valorem tariff equivalent of quotas and other kinds of tariffs. Quotas are important, for example, for Norwegian export to the EU countries. $tariff_{ivt}$ is reported as 1 + the share, and thus, $tariff_{ivt}$ = 1.01 reflects a 1% ad valorem tariff rate. For a further description of the data, see http://www.cepii.fr/CEPII/en/bdd_modele/presentation.asp?id=12.

results regarding the main variables of interest are fairly robust (see Section 3.4). I also add some other variables to control for differences at the country level: A dummy for free trade agreements (FTAs), $fta_{it,}$ to capture effects from FTAs that go beyond tariff reductions, such as reductions in non-tariff barriers;[28] GDP, gdp_{it}, and GDP per capita, $cgdp_{it}$ (both in current NOK);[29] and per capita (apparent) consumption of seafood in kg, $cons_{it}$.[30] Note that apart from the SPS variable, $tariff_{ivt}$ is the only variable varying at the product level. As is custom in the gravity literature, in the analyses, I take the natural logarithm of all continuous variables.

3.2 | Estimation method

Inserting from Equations (6) and (11) in Equations (14) – (16) allows us to express the three dependent variables as functions of different types of trade costs and demand conditions. Hence, the following reduced forms of Equations (14) – (16) are estimated:

$$\begin{aligned} \ln(y_{ivt}) = \alpha + \beta^{sps} sps_{ivt-1} + \beta^{sps*fresh} sps_{ivt-1} * fresh \\ + \beta^{tarff} \ln(tariff_{ivt-1}) + \boldsymbol{\gamma} \mathbf{x_{it}} + \mathbf{d_t} + \varepsilon_{ivt}, \end{aligned} \tag{17}$$

where y_{ivt} indicates either $export_{ivt}$, $extensive_{ivt}$ or $intensive_{ivt}$. The vector $\mathbf{x_{it}}$ represents the set of control variables that vary only in the country dimension. $\mathbf{d_t}$ is a vector of year dummies.[31]

In the main estimation model, I follow the tradition in the gravity literature of taking the natural logarithms of y_{ivt}, and then estimating (17) using OLS. Unlike most other studies, I apply a within/fixed effects estimation, due to the possible incidence of unobserved heterogeneity across countries and products that is correlated with the other explanatory variables.[32] Such heterogeneity may arise due to differences across countries as to how committed they are to report correct notifications and at a sufficiently detailed level or due to differences in demand and trading conditions at the product or country–product level not captured by the explanatory variables. For example, countries where consumers have strong preferences for a particular seafood product will have high demand for this product and may also be more concerned with the safety and quality of the product. These countries may therefore be more likely to impose an SPS on the product. In this case, unobserved consumer preferences for particular products would be positively correlated with the dependent variables as well as with the SPS variable. In the presence of such heterogeneity, estimating (17)

[28]The variable is set equal to 1 from year t and onwards if the FTA entered into force during the first half of year t and 0 otherwise. The source of the data is the Norwegian Ministry of Trade, Industry and Fisheries: http://www.regjeringen.no/en/dep/nfd/selected-topics/free-trade/partner-land.html?id=438843.

[29]Data provided by the World Bank's Governance Indicators (WGI): http://info.worldbank.org/governance/wgi/index.aspx#home.

[30]Data provided by the Food and Agriculture Organization of the United Nations (FAO): http://faostat.fao.org/site/610/default.aspx#ancor.

[31]To correct for possible product supply shocks that can vary during the period, I also did an estimation adding a dummy for the combination of years and products. Results were very similar to the ones presented in Section 3.4 both in terms of sign and significance, with one exception: The estimated $\beta^{sps*fresh}$ was insignificant in the estimation of $intensive_{ivt}$. I choose not to present these results for comparison concerns: The estimation with year-product dummies is not possible in the ppml estimation, as it does not converge due to the large number of dummies included. However, we should bear in mind the lack of robustness in the $intensive_{ivt}$ estimation as also other robustness checks shows similar results.

[32]Also, Buono and Lalanne (2012) use this method. The fixed effects are at the country–product level, which implies that I cannot estimate coefficients for time-invariant country and product variables (including dummies), as these will be soaked up by the fixed effects.

using pooled OLS or random effects would yield biased coefficient estimates. I perform a sensitivity test which indicates correlation between unobserved heterogeneity and several of the independent variables, necessitating the within/fixed effects estimation method (see Section 3.4). This method implies assuming that the error term consists of two terms, $\varepsilon_{ivt} = e_{iv} + u_{ivt}$. e_{iv} captures time-invariant differences between countries and products, and it is estimated implicitly in the model (i.e., the method is equivalent to including country–product dummies). It should be emphasised that using the within/fixed effects estimation method implies looking for an impact on export of a given product to a given country over time. Hence, I do not investigate whether exports differ to countries that have imposed an SPS on a particular product, but rather whether, on average, export of a particular product to a particular country changes when the country imposes an SPS. Cross-sectional heteroscedasticity is a common problem in trade data (see, e.g., Flam & Nordström, 2011). I therefore compute cluster robust standard errors at the country level.[33]

sps_{ivt} and $tariff_{ivt}$ may suffer from contemporaneous endogeneity: a positive shock to export may induce a country to impose an SPS or to raise tariffs. There might also be a problem with violation of the strict exogeneity assumption underlying the within/fixed effects model. This assumption requires the u_{ivt}s to be uncorrelated with the explanatory variables in *any* year (not only year t). Lagging the variables solves the first problem and mitigates the other: as pointed out by Wooldridge (2012), given that contemporaneous exogeneity holds, violation of the strict exogeneity assumption is often less severe in the within/fixed effects model when the total number of time periods, T, is large. In this analysis, $T = 10$, which is fairly large in this context. This is the reason for using lagged values of the two variables in Equation (17).[34]

Using the natural logarithm of $export_{ivt}$, $extensive_{ivt}$ or $intensive_{ivt}$ as dependent variables in the main estimation forces us to drop all observations where Norwegian export of a particular product to a particular country is zero. Therefore, the main model may suffer from sample selection bias. In a sensitivity analysis, I apply a panel version of the pseudo-Poisson maximum likelihood method (PPML) proposed by Santos Silva and Tenreyro (2006) to allow for the inclusion of the observations indicating zero export. The analysis indicates that the main results are fairly robust to the exclusion of these observations (see Section 3.4).

3.3 | Sample characteristics

The sample includes WTO members only, as these are the only countries for which there are data for SPS notifications. Seven countries joined the WTO during the sample period and are therefore not included in all years. As a consequence of non-membership in the WTO, observations amounting to approximately 13% of the total Norwegian seafood export value throughout the period have been dropped. I also drop observations covering an additional 22% of total export value due to lack of tariff data. Note, however, that results for sps_{ivt-1} are fairly robust to including these observations and at the same time dropping the tariff variable (see Section 3.4). After these adjustments,

[33]This allows for interdependence of intracountry errors such as serial correlation. Cameron and Miller (2011) suggest several guidelines for determining the appropriate level of clustering. First, if one is interested in the estimated coefficient for an aggregated explanatory variable, one should cluster at the level of aggregation of that variable. The main explanatory variables of interest are the SPS variables. As explained above, most of these are reported at the four-digit HS level for each country. Second, clusters should be implemented at the most aggregate level where intracluster serial correlation is likely to occur. By clustering at the country level, we account for serial correlation both within countries and within country–product groups over time. See also Angrist and Pischke (2008).

[34]fta_{it} is not lagged due to the way it is constructed. See footnote 28.

the sample consists of an unbalanced panel with 104 countries and nine products. It covers 64% of total export value. In the main analysis, all observations indicating zero export drop out of the analysis, and the sample consists of 3,693 observations distributed on 591 country–product groups. 58% of the observations are covered by an SPS, and 28% of observations are for fresh products. A slightly larger per cent of fresh product observations is subject to an SPS than frozen-product ones (60.7 vs. 57.0). By contrast, in the PPML analysis, only intragroup observations that are zero for all years are dropped, resulting in an increase in the number of observations included by more than one third.

3.4 | Results

3.4.1 | The main model

Results from the main model are shown in Table 1. The main explanatory variable of interest is sps_{ivt-1}. I first perform an analysis without the interaction with *fresh*. This analysis clearly demonstrates negative impacts from sps_{ivt-1} on all dependent variables. If an importing country imposes an SPS on a given product, total Norwegian export of the product to the country is reduced by 29%.[35] This reduction is due to the reduced number exporters as well as reduced exports per firm (although the latter effect is only significant at the 10% level). Consequently, there seems to be an overall dominating cost effect of SPSs on Norwegian seafood export.

Second, I investigate whether there is a heterogeneous effect across different product types by including the interaction term as demonstrated in Equation (17). The estimated β^{sps} now show the effect from SPSs on export of non-fresh seafood, whereas that from $\beta^{sps*fresh}$ show how the effect change for fresh seafood. The *net* effect of SPSs on fresh seafood is given by $\beta^{sps} + \beta^{sps*fresh}$. The table clearly demonstrates that there are large differences across products. The estimated effects from SPSs are negative and significant for non-fresh seafood for all three dependent variables. Thus, the results confirm those from the analysis without the interaction term: There seems to be a dominating cost effect from SPSs for non-fresh products. However, as shown by the positive and significant estimate of $\beta^{sps*fresh}$, this cost effect is cancelled out by a counteracting trade-enhancing effect for fresh products. A plausible explanation for the result in line with the model presented in Section 2 is that SPSs serve to reduce consumer uncertainty about the quality and safety of fresh seafood products and thereby increase demand. However, even though the estimated coefficients for the interaction terms are significantly positive in all three models, the effect is only strong enough to produce a significantly positive *net* effect of exporting fresh products for the intensive margin and then only at 10% level. Thus, for total export and the extensive margin, the demand effect appears to be strong enough to cancel out, but not to dominate the cost effect.[36,37]

The estimated effects are large: the decrease in total export for a non-fresh product associated with a country imposing an SPS is 47%. The net increase for a fresh product is 58%, but as mentioned, this net effect is not significantly different from zero.

[35]To find the estimated percentage change in y_{ivt}, $\%\hat{\Delta} y_{ivt}$, associated with sps_{ivt-1} changing from 0 to 1, we must use the following formula: $\%\hat{\Delta} y_{ivt} = 100 * (\exp \hat{\beta}^{sps} - 1)$, where $\hat{\beta}^{sps}$ is the estimated β^{sps} (Wooldridge, 2003).

[36]The p-values for the net effects (i.e., the estimated $\beta^{sps} + \beta^{sps*fresh}$) are, respectively, 0.10, 0.71 and 0.06 for total export, the extensive margin and the intensive margin (not shown in the table).

[37]I also tried running separate regressions for non-fresh and fresh products. In the former, the effect from SPS was negative and significant, whereas in the latter, the effect was not significant. These results are in line with the results from the regressions presented.

TABLE 1 Results from the within/fixed effects OLS model

Variable	ln($export_{ivt}$)		ln($export_{ivt}$)		ln($extensive_{ivt}$)		ln($extensive_{ivt}$)		ln($intensive_{ivt}$)		ln($intensive_{ivt}$)	
	Coef.	SE	Coef.	SE	Coef.	SE	Coef.	SE	Coef.	SE	Coef.	SE
sps_{ivt-1}	−0.347**	0.148	−0.626***	0.172	−0.178**	0.075	−0.254***	0.081	−0.170*	0.097	−0.372***	0.113
sps_{ivt-1}*fresh			1.087***	0.306			0.298**	0.118			0.789***	0.246
ln($tariff_{ivt-1}$)	−1.162*	0.624	−1.118*	0.602	−0.768***	0.179	−0.755***	0.176	−0.395	0.556	−0.363	0.541
fta_{it}	0.531**	0.219	0.554**	0.237	0.176**	0.085	0.182**	0.087	0.356**	0.168	0.372**	0.182
ln(gdp_{it})	2.442***	0.705	2.457***	0.720	0.909*	0.526	0.913*	0.531	1.533***	0.278	1.544***	0.283
ln($cgdp_{it}$)	−1.361*	0.713	−1.394*	0.729	−0.598	0.543	−0.607	0.548	−0.763***	0.287	−0.787***	0.292
ln($cons_{it}$)	0.127	0.326	0.164	0.316	0.313*	0.160	0.323**	0.157	−0.186	0.215	−0.159	0.209
Constant	−38.56***	12.19	−38.77***	12.48	−17.88**	8.812	−17.93**	8.885	−20.68***	0.061	−20.84***	5.575
No. of obs	3,693		3,693		3,693		3,693		3,693		3,693	
R^2	0.12		0.13		0.13		0.13		0.09		0.10	
No. of groups	591		591		591		591		591		591	

Notes: $export_{ivt}$ = export value of product *v* to country *i*, $extensive_{ivt}$ = number of firms exporting product *v* to country *i*, $intensive_{ivt}$ = their average export value. Estimation method: within/fixed effects OLS (fixed effects at the country–product level). Year dummies included, but not reported. Standard errors in parentheses (cluster robust errors with clustering at country level).
*, ** and ***significance level at 10%, 5% and 1%, respectively.

Regarding the other trade cost variables, increased tariffs have a negative influence on total exports and the extensive margin, but there is no indication of an influence on the intensive margin. This is consistent with the Chaney (2008) model with Pareto-distributed firm productivity (Lawless, 2010), or with the trade model in Medin (2017), where initially equal firms split into exporters and non-exporters due to fixed export costs in combination with "Armington" preferences. The effect of tariff reductions on the extensive and intensive margins of French exporters for years 1993–2002 was investigated in Buono and Lalanne (2012). They used a within/fixed effects estimation as here and found that increased tariffs were associated with a decrease in the intensive margin, but not the extensive one. Hence, they concluded that tariff reductions following the Uruguay Round did not help new firms starting to export, and they only helped incumbent exporters to export more. Here, I find exactly the opposite—reduced tariffs can indeed be an important policy tool for encouraging firms to export to new markets.

Free trade agreements seem to have a positive influence beyond the tariff-reducing effect, as the estimated coefficients for fta_{it} are positive for all dependent variables. This may be due to reductions in non-tariff barriers caused by the FTAs. Other studies vary as regards the effects of FTAs on trade. Kohl (2013) presents a survey of empirical studies using gravity analysis and finds no effect from FTAs on trade in almost 65% of the cases investigated. Moreover, only about one quarter of the agreements lead to increased trade, and 10% actually had a negative impact.

As to the remaining controls, GDP is found to influence all margins of trade positively, as expected. GDP per capita has a clear negative influence on $intensive_{ivt}$ and a weakly significant negative influence on $export_{ivt}$. Consequently, after correcting for total income, wealthier countries seem to demand less seafood. This can be due to seafood being a necessity good. Per capita consumption of seafood has a positive effect on $extensive_{ivt}$. Within R^2 is quite low, ranging from 4% to 10%. However, R^2 is expected to be low in gravity analyses at such a disaggregated level.

3.4.2 | Robustness checks

In order to test the robustness of the results, especially regarding the impact of SPSs, I have run some alternative estimations of (17). First, I test for sample selection bias by investigating how sensitive the results from the main model are to including the observations indicating zero export (see Section 3.2). In the panel PPML method applied for this purpose, I use levels instead of logs of the dependent variables. The model therefore includes the zero observations in a natural way. I use within/fixed effects estimation to correct for time-invariant unobserved heterogeneity at the country–product level, just as in the main model.[38] The results are shown in Table 2. Results for sps_{ivt-1} and its interaction with *fresh* are fairly similar to those from the main model, in terms of both sign and significance.[39] The main difference is that the positive effects for fresh seafood products are somewhat larger for $export_{ivt}$ and $intensive_{ivt}$. Also, effects of $tariff_{ivt-1}$, fta_{it} and gdp_{it} are similar in the two models, whereas effects from per capita GDP and seafood consumption vary.

[38]In the PPML estimation, we do not cluster at the country level, as in the main analysis, but rather at the country–product level, because the applied software (Stata 13) allows only for clustering at the same level as the fixed effects in this type of estimation.

[39]In the PPML estimation, $E(y_{ivt}|\mathbf{z_{ivt}}) = \exp(\alpha + \boldsymbol{\beta}\mathbf{z_{ivt}})$, where $\mathbf{z_{ivt}}$ is a vector of the independent variables and $\boldsymbol{\beta}$ is the coefficient vector. Consequently, the βs can be interpreted in the same way as in the OLS estimation (i.e., when a z_{ivt} is given in log, the corresponding β roughly reflects the elasticity, just as in the log-log OLS model; Wooldridge, 2012).

TABLE 2 Results from within/fixed effects pseudo-Poisson maximum likelihood regression

	ln($export_{ivt}$)		ln($extensive_{ivt}$)		ln($intensive_{ivt}$)	
Variable	**Coef.**	***SE***	**Coef.**	***SE***	**Coef.**	***SE***
sps_{ivt-1}	−0.471***	0.153	−0.256***	0.055	−0.374***	0.091
sps_{ivt-1}*fresh*	1.271***	0.324	0.301***	0.101	1.025***	0.210
ln($tariff_{ivt-1}$)	−1.532**	0.763	−0.802***	0.158	−0.451	0.465
fta_{it}	0.431***	0.113	0.265***	0.071	0.230*	0.129
ln(gdp_{it})	3.246**	1.521	0.897***	0.231	1.625*	0.837
ln($cgdp_{it}$)	−2.604	1.620	−0.537**	0.236	−1.200	0.828
ln($cons_{it}$)	0.998**	0.407	0.358***	0.109	0.161	0.212
No. of obs	5,751		5,751		5,751	
Log pseudolikelihood	−15,094,969,294		−7,157		−1,697,534,232	
No. of groups	591		591		591	

Notes: $export_{ivt}$ = export value of product *v* to country *i*, $extensive_{ivt}$ = number of firms exporting product *v* to country *i*, $intensive_{ivt}$ = their average export value. Estimation method: within/fixed effects Poisson maximum likelihood (fixed effects at the country–product level). Year dummies included, but not reported. Standard errors in parentheses (cluster robust errors with clustering at country–product level).

*, ** and ***significance level at 10%, 5% and 1%, respectively.

These results indicate that sample selection bias due to the omission of the zero observations in the main model is not severe, despite the large number of such observations. Including the zero trade flows do not alter the conclusions regarding the main variables of interest from the main analysis. Also, other scholars have found small selection bias in gravity models (Helpman et al., 2008; Santos Silva & Tenreyro, 2006).

Second, I check whether three alternative treatments of the tariff variable influence the results from the main analysis. In the first, I do not fill in the missing values of $tariff_{ivt-1}$ by interpolation but rather by setting them equal to the closest preceding observation.[40] In the second, I use the same tariff variable as in the main analysis, but estimate (17) for years 2002–07 only. This means I rely on interpolation between the observed values, as before, but abstain from guessing the values for years succeeding that with the latest observation. In the third, I estimate (17) without the tariff variable. This leads to a large increase in the number of products included: from 9 to 102. Also, the number of countries covered by the sample increases—from 110 to 119. (See Section 3.1. for details on the construction of $tariff_{ivt-1}$.) Table 3 shows the results for the SPS variables and $tariff_{ivt-1}$ from these three analyses. As demonstrated, results for sps_{ivt-1} in the estimations of $export_{ivt}$ and $extensive_{ivt}$ are similar to those in the main analysis. The size of the coefficients varies somewhat, but in none of the estimations are the conclusions from the main analysis regarding the SPS variables altered. In the analysis without tariffs, the estimated coefficient for sps_{ivt-1} is insignificant in the estimation of $intensive_{ivt}$, thus results for this variable are somewhat robust. Results for $tariff_{ivt-1}$ are similar across the estimations where it is included.

Third, I check whether it would be adequate to use a pooled or a random-effects estimation of (17). As opposed to the within/fixed effects estimation used in the main model, in estimations

[40]That is, values in years 2002 and 2003 are set equal to the 2001 values, those for years 2005 and 2006 to the 2004 values and so on.

TABLE 3 Results for the main variables of interest and tariffs in estimations with different treatments of the tariff variable

	ln($export_{ivt}$)		ln($extensive_{ivt}$)		ln($intensive_{ivt}$)	
	Coef.	*SE*	Coef.	*SE*	Coef.	*SE*
Tariff = preceding						
sps_{ivt-1}	−0.613***	0.171	−0.246***	0.080	−0.368***	0.113
sps_{ivt-1}*fresh*	1.091***	0.306	0.301**	0.117	0.790***	0.246
ln($tariff_{ivt-1}$)	−1.028*	0.519	−0.679***	0.136	−0.350	0.477
Years 2002–07						
sps_{ivt-1}	−0.800***	0.179	−0.314***	0.089	−0.486***	0.123
sps_{ivt-1}*fresh*	1.063***	0.339	0.393***	0.144	0.670**	0.263
ln($tariff_{ivt-1}$)	−0.822	0.755	−0.493**	0.195	−0.330	0.614
No tariff variable						
sps_{ivt-1}	−0.257***	0.096	−0.136***	0.034	−0.121	0.078
sps_{ivt-1}*fresh*	0.766***	0.207	0.152**	0.065	0.614***	0.220
ln($tariff_{ivt-1}$)	n.a.		n.a.		n.a.	

Notes: $export_{ivt}$ = export value of product v to country i, $extensive_{ivt}$ = number of firms exporting product v to country i, $intensive_{ivt}$ = their average export value. Estimation method: within/fixed effects OLS (fixed effects at the country–product level). Year dummies included, but not reported. Standard errors in parentheses (cluster robust errors with clustering at country level).
*, ** and ***significance level at 10%, 5% and 1%, respectively. Only results for the main variables of interest and tariffs are shown.

like that, the coefficients would reflect differences in SPSs across countries as well as within countries over time.[41] However, in order for estimates to be unbiased, the unobserved heterogeneity term (e_{iv}) must be uncorrelated with the other explanatory variables (see Section 3.2). To test whether this holds, I add the time-invariant means of all the independent variables to Equation (17) and then estimate it using random effect as in Mundlak (1978; also see Wooldridge, 2012). I also include four time-invariant control variables at the country level, because variables like that are no longer soaked up by the fixed effects. The variables are as follows: a dummy for landlocked countries ($landlocked_i$), distance from Norway ($dist_i$), average distance to the rest of the world ($remoteness_i$) and a dummy for Nordic countries ($nordic_i$).[42] Results are shown in Table 4. The table clearly demonstrates that several of the estimated coefficients for the time-invariant means are significant, indicating that there is indeed correlation between unobserved heterogeneity and the independent variables.[43] Consequently, a pooled or random-effects approach would not be adequate.

Summing up, the results from the robustness checks indicate fairly robust results for SPS variables, trade cost variables and GDP. Results regarding GDP per capita and seafood

[41]The coefficients from a random-effects estimation are weighted averages between those from a within estimation and those from a between estimation.

[42]$remoteness_i$ is defined as $\frac{1}{n}\sum_{j=1}^{n} d_{ij}$, where d_{ji} is distance from country i to country j, and n is the number of countries. Internal distance d_{jj} is set equal to the square root of the country's area multiplied by about 0.4. Data are taken from the CEPII database *dist_cepii* (Mayer & Zignago, 2011). For $dist_i$ and $remoteness_i$, I use the Great Circle distance measured in kilometres between largest cities (the $dist_i$ variable). $dist_i$ and $remoteness_i$ are given in logs.

[43]Similar results are found in other studies (see, e.g., Egger, 2000).

TABLE 4 Testing for correlation between unobserved heterogeneity and the explanatory variables

	ln($export_{ivt}$)		ln($extensive_{ivt}$)		ln($intensive_{ivt}$)	
Variable	**Coef.**	***SE***	**Coef.**	***SE***	**Coef.**	***SE***
sps_{ivt-1}	−0.631***	0.170	−0.251***	0.079	−0.382***	0.112
Mean sps_{iv}	0.169	0.531	0.228	0.193	−0.062	0.378
sps_{ivt-1}*$fresh$	1.085***	0.303	0.301***	0.116	0.784***	0.245
Mean sps_{iv}*$fresh$	1.831***	0.639	0.305	0.214	1.511***	0.492
ln($tariff_{ivt-1}$)	−1.142*	0.597	−0.757***	0.173	−0.392	0.539
Mean ln($tariff_{iv}$)	−1.719	1.584	−0.107	0.525	−1.609	1.300
fta_{it}	0.543**	0.236	0.177**	0.086	0.361**	0.181
Mean fta_i	−1.148*	0.646	−0.489**	0.242	−0.633	0.446
ln(gdp_{it})	2.399***	0.708	0.895*	0.521	1.483***	0.279
Mean ln(gdp_i)	−1.883***	0.712	−0.679	0.521	−1.180***	0.287
ln($cgdp_{it}$)	−1.334*	0.716	−0.591	0.539	−0.721**	0.284
Mean ln($cgdp_i$)	1.096	0.777	0.512	0.540	0.561	0.344
ln($cons_{it}$)	0.162	0.308	0.318**	0.150	−0.157	0.203
Mean ln($cons_{it}$)	0.607	0.378	−0.143	0.168	0.751***	0.258
$remoteness_i$	0.408	1.670	0.744	0.685	−0.328	1.128
ln($dist_i$)	−1.526***	0.324	−0.834***	0.155	−0.690***	0.205
$landlocked_i$	−0.949**	0.474	−0.436**	0.195	−0.523	0.321
$nordic_i$	−1.188**	0.508	−0.429	0.275	−0.776***	0.272
Constant	9.700	13.839	−3.707	5.327	13.239	9.625
No. of obs	3,693		3,693		3,693	
R^2	0.36		0.41		0.31	
No. of groups	591		591		591	

Notes: $export_{ivt}$ = export value of product v to country i, $extensive_{ivt}$ = number of firms exporting product v to country i, $intensive_{ivt}$ = their average export value. Estimation method: random effects (random effects at the country–product level). Year dummies included, but not reported. Standard errors in parentheses (cluster robust errors with clustering at country level).
*, ** and ***significance level at 10%, 5% and 1%, respectively.

consumption are less robust for all dependent variables. None of the robustness checks alter the main results from this article—that there is a negative effect from SPSs on export of seafood products in general, but for fresh seafood, this is cancelled out by counteracting positive effect.

4 | CONCLUSIONS

In research on international food standards, several recent studies have indicated that the standards' impact on trade differs across products. Food standards may act as barriers to trade, but they may also lead to increased trade. Here, I have presented a theoretical explanation of this "dual face of food standards" hypothesis by slightly modifying existing models. I have also offered new empirical evidence on the impact of foreign food standards for Norwegian seafood export.

There are two explanations suggested for a positive relationship between food standards and trade. First, compliance with standards may lead to a direct increase in demand, as they may entail quality upgrades or reduce consumer uncertainty about quality and safety of products. Second, standards may distort trade and improve the competitive advantage of countries that have better capacity to comply with them.

In the empirical analysis, I employed data for Norwegian firm-level seafood export of different products to different countries. I investigated how foreign food standards influenced three variables: total export, the number of exporters (the extensive margin) and their average export (the intensive margin). WTO notification data on SPSs were used as indicators of food standards. On the whole, SPSs are shown to have a negative impact on exports. However, the impact is heterogeneous across products, and for export of fresh seafood, the negative effect is cancelled out by a counteracting positive effect. These results are in line with some other studies indicating that the impact of food standards is heterogeneous across products. Further research should be conducted to trace the exact mechanisms underlying this apparently contradictory evidence. Another interesting point for future research is how firm survival over time is affected by standards and non-tariff barriers.

ACKNOWLEDGEMENTS

I am very grateful to Arne Melchior for extensive commenting and useful inputs during the whole research process. In addition to referees, I also thank the following people for comments: Pol Antras, Ragnhild Balsvik, Andrew Bernard, Elena Besedina, Esther Ann Bøler, Thomas Von Brasch, Taran Fæhn, Rikard Forslid, Tommy Sveen, Mads Greaker and Kjetil Telle. Finally, I thank Marcus Gjems Theie for data preparation assistance and Susan Høivik for excellent copy-editing.

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How to cite this article: Medin H. Trade barriers or trade facilitators? The heterogeneous impact of food standards in international trade. *World Econ*. 2019;42:1057–1076. https://doi.org/10.1111/twec.12756