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Explaining the Size Differences of Exporter Premia: Theory and Evidence

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Abstract

Two decades of research have established pronounced exporter productivity premia (EPP) and exporter size premia (ESP). Yet, we do not know why such exporter premia differ so widely in magnitude across countries or sectors? We take this question to the theory and to the data. We derive the sectoral EPP and ESP in a standard heterogeneous firms trade model and apply the insights from the model to guide our empirical investigation of detailed Danish firm-level data. We show that a significant share of the observed variation in EPP and ESP across sectors can be accounted for by sector differences in the underlying variation in productivity dispersion, variable trade costs, the ratio of fixed export costs to fixed costs of production, and the elasticity of demand.

JEL: F12, F14 , D24, O47

Keywords: Intra-industry trade, exporter productivity, exporter size premium, firm-level data, heterogeneous firms.

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1 Introduction

The existence of persistent productivity dispersion is a vivid research topic in fields ranging from macroeconomics to industrial organization. Maybe the most prominent manifestation of such productivity dispersions stems from the field of international economics, namely the stylized fact of positive exporter premia in terms of productivity or size and sales.¹ Combined with elegant new theories (e.g. Melitz, 2003) exporter productivity premia (EPP) – as well as the closely related exporter firm size or sales premia (ESP) – have intrigued empirical and theoretical researchers in international economics for the past decades; see Greenaway and Kneller (2007), Wagner (2007, 2012), Redding (2011) and Bernard et al. (2012) for surveys. Most recently Mrázová et al. (2015) have been able to map fundamental theoretical conditions under which productivity and sales distributions are linked, this being of particular importance since sales distributions are much easier to observe.

Despite the impressive volume of analyses, the central question as to the underlying drivers of the EPP and ESP magnitude remains in fact open. Albeit, as has been pointed out by, *inter alia*, Syverson (2011), an understanding of the underlying drivers of the dispersion observed in the data is essential for many policy areas such as competition policy or government measures aiming at promoting productivity growth in the economy. In particular, the understanding of the EPP and ESP developed in this paper is an important stepping stone when thinking about effects and effectiveness of trade policies or export promotion policies. For example, larger EPP make it *ceteris paribus* more attractive in terms of production efficiency to design industrial policies that reallocate production resources (market shares) from non-exporters towards incumbent exporters. Viewed through this lens, optimal industrial policies (say export guarantees vs. general production subsidies) would in fact depend on the magnitude of the EPP.²

What determines the size of such premia and why is it that exporter premia vary so widely in magnitude? Empirically, exporter productivity premia range, for example in Europe, from 7 (0)% in Sweden to 58 (10)% in Belgium for identical pooled (fixed effects) estimation specifications on comparable data (ISGEP, 2008, Table 4). At the sectoral level – the perspective taken in the present paper – the issue is even more pronounced.³ For Den-

¹We use the term premia in line with previous literature referring to differences between firms that export and firms that do not.

²On a different note, our work also highlights that certain policies (such as policies that reduce trade costs for firms) can impact on the size of the EPP and ESP.

³We obtain EPP and ESP estimates from industry-by-industry regressions of TFP respectively sales on an exporter status dummy conditioning on firm-specific and time-

mark total factor productivity based fixed effects EPP estimates across 89 manufacturing sectors span from -38% to 467%. Similarly, the exporter sales premia ESP estimates span from -22% to 278%. On first sight, theory is silent on such differences, thus posing a challenge for the interpretation of empirical results.⁴

To be fair, short of classifying EPP and ESP differences in magnitude simply as idiosyncratic characteristics of sectors or differences in econometric estimation strategies, a number of renown empirical papers have elaborated on the issue. Typically, papers that report differences in EPP and ESP magnitude have addressed the issue by conducting additional investigation, for example by splitting samples and by providing plausible rationalizations; see e.g. Aw and Hwang (1995), Farinas and Martin-Marcos (2007), or Merino (2004). Similarly, ISGEP (2008) establishes and discusses intensively the identified size differences between country EPP, and Bellone et al. (2014) report larger EPP for Japan than for France.⁵ Recently Powell and Wagner (2014) took a look at the issue of exporter premia size differences by examining the EPP along the productivity distribution via quantile regression. The current paper takes a different direction.

We explore the link between a stylized heterogeneous firms trade model and the empirical patterns of EPP and ESP magnitude. In doing so, the paper contributes on three frontiers. First, we provide a theoretical underpinning of why and how sectoral EPP and ESP differ as driven by sector-specific characteristics in an adapted Melitz (2003) model. Previous theoretical literature has – despite of deriving productivity rankings of exporters and non-exporters – not examined the actual premia, i.e. the theoretical EPP or ESP. Second, in our empirical investigation, we construct (i.e. by structural estimation) the deep parameters pinpointed by the model. Obviously, the downside of guiding the empirical investigation within one specific model of firm heterogeneity is that it can not capture all the empirical variation known to matter in the data. Still, we find considerable heterogeneity in the deep parameters across industries, which in turn results in substantial differences in EPP and ESP. For example, in our data, once the underlying

specific fixed effects (see Section 5).

⁴Note, that negative exporter premia are not quite as exceptional as one would suspect, see Schröder and Sørensen (2012) for a theoretical underpinning and examples of empirical findings.

⁵In contrast to both these papers, much of the remaining literature on country-level exporter premia are single-country studies (see the surveys of Greenaway and Kneller, 2007; Wagner, 2007, 2012; Bernard et al., 2012; and Hayakawa et al., 2012). In particular, the implementation of different estimation strategies makes comparisons across studies difficult.

industry characteristics suggested as the drivers of EPP and ESP magnitude by our theory are taken into consideration, then 43% of the empirically observed cross-sectoral variation in TFP based EPP differences is accounted for. Thirdly, the paper is able to identify new and so far unreported empirical regularities, such as the effects from the underlying variation in productivity dispersion on the EPP and ESP. In sum, our findings offer first explanations for the substantial size differences in EPP and ESP that have been reported in the literature for over two decades.

The remainder of the paper is structured as follows. Section 2 introduces and calculates the sectoral EPP and ESP in a two-country symmetric multi-sector Melitz (2003) model augmented with Pareto-distributed productivities. We find that within the model larger sectoral EPP and ESP are linked to a larger degree of heterogeneity in the productivity distribution of a sector, higher trade costs (both fixed and variable), and for the ESP also to a larger within-sector elasticity of demand. Section 3 presents our data and empirical set-up. Based on 18 years of firm-level data for the universe of Danish firms, we provide structural estimates of the deep parameters identified in the theoretical model and subsequently we examine firm-level EPP and ESP estimations across 3-digit level industries. Our results are shown in Section 4 and document – in line with the model – the role of the underlying heterogeneity in productivity and the size of fixed export costs (measured relative to domestic production fixed costs) for the ESP and EPP. Moreover, we find a negative relation between the sectoral elasticity of demand and both exporter premia and cannot identify a clear relation between variable trade costs and the EPP and ESP. Section 5 discusses future empirical research and demonstrates that a sizable share of the sectoral EPP and ESP variation found in our data can be accounted for by the underlying industry characteristics suggested by our analysis. Section 6 summarizes.

2 Exporter Premia in Theory

Existing theoretical work provides the well-known rankings of firms according to marginal productivity or size. Such a ranking implies that firms above the export threshold export and firms below do not export. While a generally positive sign for the EPP and ESP follows from this exercise, inference on the drivers of premia magnitude is not provided. We resolve this issue and present insights from theory addressing the relation of the deep parameters of the model to the sectoral EPP and ESP. A simple thought experiment illustrates the task. Consider a standard heterogeneous firms trade model and add one extremely productive exporting firm. At first sight, this should increase

the EPP and ESP. Yet, empirical measures of exporter premia compare the group of exporters to non-exporters. Thus, via general equilibrium effects the addition of a highly productive firm toughens the exit and exporter cut-off and therefore changes the composition of firms in both groups. Hence, it becomes a non-trivial task to compute the overall effect on the EPP and ESP.

Against this backdrop, we derive the exporter premia (EPP and ESP) in a symmetric two-country variant of Melitz (2003). Thus, we bring the model closer to what is actually examined in empirical work. In line with the literature we only consider steady states. We make two modifications to the workhorse model. First, in order to reflect cross-sectoral variation in EPP or ESP we rewrite the model to include J heterogeneous sectors. Second, to ensure tractability and transparency and to obtain clear-cut analytical predictions on how the premia depend on industry-specific characteristics, we adopt the conventional assumption of productivities being Pareto-distributed.⁶ Accordingly, all the well-known properties of the Melitz model are preserved in what follows.

2.1 Households

A representative household supplies L units of labour inelastically to the labour market and derives utility from consumption of different varieties from the J sectors. The utility function of the household reads

$$U = u(C_1, C_2, \dots, C_J),$$

where C_j denotes the consumption of the sector j specific composite consumption bundle.⁷ The composite consumption bundles are of the Dixit-Stiglitz (1977) CES type

$$C_j = \left(\int_{\omega \in \Omega_j} (c_j(\omega))^{\frac{\sigma_j-1}{\sigma_j}} d\omega \right)^{\frac{\sigma_j}{\sigma_j-1}} \quad \text{for } j = 1, 2, \dots, J, \quad (1)$$

where $c_j(\omega)$ denotes consumption of variety ω of sector j , Ω_j is the endogenously determined set of varieties (both domestic and foreign) from sector j

⁶The application of the Pareto distribution within heterogeneous firms trade models is by now common. Moreover, the Pareto approximates the distribution of productivity of firms found in empirical work (see e.g. Luttmer, 2007) and resembles the empirical patterns we find for Danish firms at least for the right tail of the distribution, see Figure 1 and Section 3.1.

⁷We assume that all first-order partial derivatives are positive, i.e. utility increases in all the sector-specific consumption bundles (defined by (1)).

available to the household and σ_j is the elasticity of substitution between any two varieties within sector j (equivalently the elasticity of demand). Demand for variety ω of sector j becomes

$$c_j(\omega) = C_j \left(\frac{p_j(\omega)}{P_j} \right)^{-\sigma_j} \quad \text{for } j = 1, 2, \dots, J \text{ and } \forall \omega \in \Omega_j, \quad (2)$$

where $p_j(\omega)$ is the price of variety ω of sector j and P_j is the price of buying one unit of the composite bundle C_j .⁸

2.2 Firms

All J sectors feature monopolistic competition. Accordingly, firms take the sector-level variables C_j and P_j as given. Each firm produces one unique variety ω . Labour is the only input factor and remunerated at the economy wide wage rate w . Entry into a sector, i.e. inventing a new variety, is associated with sunk costs of $f_{j,e}$ units of labour. The creation of a new variety entails a variety-specific and fixed productivity, $\varphi_j(\omega)$.⁹ Since firms within a sector only differ in terms of productivity, we omit unnecessary notation and simply refer to φ in the reminder of the paper. Realized variety-specific productivity is stochastic due to the random nature of research and development (R&D) processes. We assume that productivities are drawn from known sector-specific Pareto distributions with location parameters $\varphi_{0,j}$ and shape parameters k_j .¹⁰ The cumulative distribution function thus reads

$$G_j(\varphi) = 1 - \left(\frac{\varphi}{\varphi_{j,0}} \right)^{-k_j} \quad \text{for } \varphi > \varphi_{j,0} \text{ and } j = 1, 2, \dots, J. \quad (3)$$

Firms from a competitive fringe enter the industry until expected profits from entry equal the sunk costs of entry. Firms that decide to produce face fixed

⁸The price index reads

$$P_j = \left(\int_{\omega \in \Omega_j} (p_j(\omega))^{1-\sigma_j} d\omega \right)^{\frac{1}{1-\sigma_j}} \quad \text{for } j = 1, 2, \dots, J.$$

⁹Staying true to these well-established features of modeling firms ignores a range of factors that recent empirical work has uncovered. Most importantly, works such as Manova and Zhang (2012) and Di Comite et al. (2014) have established that quality and demand (taste) shifts play an important role in explaining firm sales. Appendix B provides an extended version of the present model that takes account of quality heterogeneity and demand shifters.

¹⁰We impose the parameter restriction $k_j > \max(2, \sigma_j - 1)$ in order to ensure finite variance of the productivity distribution and finite expected profits prior to entry.

costs of production f_j (units of labour). Accordingly, the labour requirement of a firm with productivity φ that produces q units becomes $l_j(q|\varphi) = \frac{q}{\varphi} + f_j$. In addition, firms have the option to export. Exporting is subject to fixed export costs of $f_{j,x}$ units of labour and variable export costs of the iceberg type, i.e. firms must ship τ_j units for one unit to arrive.

Given the constant elasticity of demand, cf. (2), firms in each sector set prices as constant mark-ups on marginal costs implying that domestic (subscript d) and export market (subscript x) prices are given by

$$\begin{aligned} p_{j,d}(\varphi) &= \frac{\sigma_j}{\sigma_j - 1} \frac{w}{\varphi} \\ p_{j,x}(\varphi) &= \tau_j p_{j,d}(\varphi) . \end{aligned}$$

Sales (revenue) at home and abroad become

$$s_{j,d}(\varphi) = C_j (P_j)^{\sigma_j} \left(\frac{\sigma_j}{\sigma_j - 1} \frac{w}{\varphi} \right)^{1-\sigma_j} , \quad (4)$$

$$s_{j,x}(\varphi) = \tau_j^{1-\sigma_j} s_{j,d}(\varphi) , \quad (5)$$

and total sales are

$$s_j(\varphi) = s_{j,d}(\varphi) + I_{j,x}(\varphi) s_{j,x}(\varphi) ,$$

where $I_{j,x}(\varphi)$ is a dummy variable for export status.

Accordingly, reduced form profits in the domestic market and in the export market of a sector are given by

$$\begin{aligned} \pi_{j,d}(\varphi) &= B_j(\varphi)^{\sigma_j-1} - w f_j \\ \pi_{j,x}(\varphi) &= B_j(\varphi)^{\sigma_j-1} \tau_j^{1-\sigma_j} - w f_{j,x} \end{aligned}$$

where $B_j \equiv C_j (P_j)^{\sigma_j} \left(\frac{\sigma_j}{\sigma_j-1} \right)^{-\sigma_j} \frac{1}{\sigma_j-1} (w)^{1-\sigma_j}$ is a sector-specific demand component.

Only firms with a sufficiently high productivity find it profitable to supply a given market. Consequently, firms self-select according to productivity into exiters ($\varphi < \varphi_{j,d}^*$), pure domestic non-exporters ($\varphi_{j,d}^* \leq \varphi < \varphi_{j,x}^*$) and exporters ($\varphi \geq \varphi_{j,x}^*$) where the exit thresholds, $\varphi_{j,d}^*$, and export thresholds of a sector ($\varphi_{j,x}^*$) are defined by

$$\pi_{j,d}(\varphi_{j,d}^*) = 0 \Leftrightarrow \varphi_{j,d}^* = \left(\frac{w f_j}{B_j} \right)^{\frac{1}{\sigma_j-1}} \quad (6)$$

$$\pi_{j,x}(\varphi_{j,x}^*) = 0 \Leftrightarrow \varphi_{j,x}^* = \left(\frac{w f_{j,x} \tau_j^{\sigma_j-1}}{B_j} \right)^{\frac{1}{\sigma_j-1}} . \quad (7)$$

We impose the parameter restriction $f_{j,x}\tau_j^{\sigma_j-1} > f_j$ which ensures that firms – in line with empirical evidence – partition into exporters and non-exporters.

2.3 Theoretical EPP and ESP Predictions

Within the above specification we are able to compute the sectoral EPP and ESP. In particular, we derive theoretical exporter premia measures that compare the difference in average productivity or sales (as a measure of size) between the group of exporters and the group of non-exporters relative to the average productivity or sales of the group of non-exporters.

The exporter premia $EP_{j,z}$ in sector j are defined as:

$$EP_{j,z} \equiv \frac{\mathbb{E}[z_j(\varphi) | \varphi \geq \varphi_{j,x}^*] - \mathbb{E}[z_j(\varphi) | \varphi_{j,d}^* \leq \varphi < \varphi_{j,x}^*]}{\mathbb{E}[z_j(\varphi) | \varphi_{j,d}^* \leq \varphi < \varphi_{j,x}^*]}, \quad (8)$$

whereby $z_j(\varphi) = \{\varphi, s_j(\varphi)\}$ determines whether we compute EPP or ESP, respectively.

By using the Pareto distribution (3), one can rewrite (8) as

$$EP_{j,\varphi} = \frac{\left(\frac{\varphi_{j,d}^*}{\varphi_{j,x}^*}\right)^{-k_j} - 1}{\left(\frac{\varphi_{j,d}^*}{\varphi_{j,x}^*}\right)^{-k_j+1} - 1}, \quad (9)$$

for the EPP, and as

$$EP_{j,s} = \left(1 + \tau_j^{1-\sigma_j}\right) \frac{\left(\frac{\varphi_{j,d}^*}{\varphi_{j,x}^*}\right)^{-k_j} - 1}{\left(\frac{\varphi_{j,d}^*}{\varphi_{j,x}^*}\right)^{-k_j+(\sigma_j-1)} - 1}, \quad (10)$$

for the ESP, respectively.¹¹

Some discussion is in place: In the present stylized model there is a positive and monotone relationship between productivity and sales in a given market that is common across firms (see equations (4) and (5)).¹² In fact, in the present case both productivities and sales in a given market are Pareto-distributed due to the assumption of CES preferences (see e.g. Mrázová

¹¹See appendix A for an elaboration of (9) and (10), as well as for derivations of the results that follow.

¹²In more advanced models with quality heterogeneity and demand heterogeneity the relationship will be contingent on quality and demand. Accordingly, once the relation between productivity and sales is less clear, so will the relation between EPP and ESP. See Appendix B for such an extension of the model.

et al., 2015). Accordingly, EPP and ESP are closely related to each other in this framework, see (9) and (10). However, even in this stylized setting there are differences. The reason is that while productivities are Pareto with shape parameter k_j , sales are Pareto with shape parameter $\frac{k_j}{\sigma_j-1}$. Moreover, when firms start to export, there is a discrete upward jump in sales and the slope/curvature of the relationship changes due to the iceberg trade costs. Hence, it is important to consider both premia in the present framework.

It is important to note that given CES consumption bundles and Pareto-distributed productivity (φ), a sector's structure – and accordingly the sectoral EPP and ESP – only depends on parameters specific to the sector in question and not on the general equilibrium values of all B_j for $j = 1, 2, \dots, J$. Why is this so? In general, no clear-cut results can be provided unless some structure on the productivity distribution is assumed. The Pareto assumption evoked in the present framework is particularly convenient in that respect. A key property of the distribution is that a Pareto truncated from below (due to selection into production and exporting) remains Pareto with the same shape parameter as the non-truncated distribution. This implies that only the ratio between thresholds matters. Even though other sectors will affect the general equilibrium and thereby both thresholds, this happens such that the ratio between thresholds within a sector is unchanged. This is why the exporter productivity premium does not depend on parameters from other sectors. Thus, the Pareto enables us to conclude on the effects from a change in the underlying productivity distribution through a change in a single parameter without making the statement contingent on initial values of thresholds (and other parameters).¹³

We can now state:

Proposition 1. *The theoretical sectoral exporter productivity premium, $EP_{j,\varphi}$, and size premium, $EP_{j,s}$, in sector j only depend on the sector parameters $k_j, f_{j,x}, \tau_j, f_j$ and σ_j and are given by*

$$EP_{j,\varphi} = \frac{\left(\frac{f_{j,x}}{f_j}\right)^{\frac{k_j}{\sigma_j-1}} \tau_j^{k_j} - 1}{\left(\frac{f_{j,x}}{f_j}\right)^{\frac{k_j-1}{\sigma_j-1}} \tau_j^{k_j-1} - 1} - 1. \quad (11)$$

$$EP_{j,s} = \left(1 + \tau_j^{1-\sigma_j}\right) \frac{\left(\frac{f_{j,x}}{f_j}\right)^{\frac{k_j}{\sigma_j-1}} \tau_j^{k_j} - 1}{\left(\frac{f_{j,x}}{f_j}\right)^{\frac{k_j}{\sigma_j-1}-1} \tau_j^{k_j-(\sigma_j-1)} - 1} - 1. \quad (12)$$

¹³Reassuringly, we find strong empirical support for the Pareto distribution in the Danish firm-level data (see Figure 1 and the discussion in Section 3.1).

Proof. The premia follow from evaluating (9) and (10) using the thresholds given by (6) and (7), see Appendix A. $\square$

Proposition 1 gives a series of results on how sector parameters affect the magnitude of a sector's EPP and ESP.¹⁴

Corollary 1. *Sectors that are more heterogeneous in their productivity distribution (lower k_j , implying a higher productivity dispersion) have larger exporter productivity and size premia.*

Corollary 2. *Sectors that have higher variable trade costs (τ_j) have larger exporter productivity and size premia.*

Corollary 3. *Sectors that have higher fixed export costs to fixed cost of production ratios ($\frac{f_{j,x}}{f_j}$) have larger exporter productivity and size premia.*

Corollary 4. *Sectors that have larger sectoral elasticity of demand (σ_j) have larger exporter size premia.*

Corollary 5. *The effect of the sectoral elasticity of demand (σ_j) on the exporter productivity premia is ambiguous.*

Corollaries 1 to 4 provide clear and testable predictions on the drivers of EPP and ESP magnitude.

To understand Corollary 1, note that the degree of heterogeneity increases (higher productivity dispersion) when the right tail of the distribution has more mass (lower k_j). More mass in the tail in turn increases the density among highly productive large firms – both among exporters and among non-exporters. However, the effect is less pronounced for non-exporters because their productivity distribution, unlike that of exporters, is right-truncated. Hence, the EPP and ESP increase with the degree of heterogeneity.¹⁵

There is an important difference between corollaries 4 and 5. While we establish a clear theoretical relation between the sectoral elasticity of demand and the ESP (Corollary 4), theory does not provide a clear expected sign for the relation between the elasticity of demand and the magnitude of the EPP (Corollary 5). However, in the parameter range, where $f_{j,x} < f_j$, one

¹⁴The proofs of the Corollaries follow from the partial derivatives of the expression for the sector-specific premia (11) and (12), see Appendix A.

¹⁵The observation that the dispersion of the productivity distribution is an important fundamental is also made by Helpman et al. (2004) in the context of export channel choices, i.e. their investigation of the relation between direct exports and foreign affiliate sales in various destination markets.

can show that a larger σ_j is associated with larger EPP, and vice versa for $f_{j,x} > f_j$ (see Appendix A).

A final remark is in order: The productivity measure, and accordingly the premium measure $EP_{j,\varphi}$ developed above, is in the literature following Melitz (2003) the theoretical counterpart to empirical observable productivity measures such as TFP or value-added per employee. However, as shown in Schröder and Sørensen (2012) there is a settle difference between theoretical and empirical measures. The term φ from Melitz (2003) measures marginal production efficiency and thus captures the increase in output (measured in physical units) per extra worker, i.e. a marginal productivity measure in terms of output units. Empirical measures are in contrast based on monetary values (not physical units) and include fixed costs of production, i.e. the empirically observable EPP measures include pricing decisions and are based on average costs. In Appendix C we show that the theoretical EPP in (9) above is not an accurate representation of the empirically observable EPP. In contrast, the sales-based ESP measure from above is already stated in monetary values, thus the theory-based ESP measure is fully in line with actual empirical ESP observations. This underlines once more the relevance of using size measures such as ESP alongside the common EPP measures.

3 Data and Empirical Estimation Strategies

In line with the foundation outlined in Section 2, our empirical focus lies on identifying productivity and size differentials between exporters and non-exporters. Here our approach is to match moments from the model with moments from data to obtain industry-level structural estimates of the parameters shown in Section 2 to theoretically affect EPP and ESP. Subsequently, we include these structurally estimated parameters as interaction terms in simple firm-fixed effects models for log sales and two different productivity measures. The core of the analysis then examines how the empirically observed EPP and ESP correlate with the estimated sector-level structural parameters.

Wagner (2007, 2012) provides comprehensive surveys of the empirical literature and finds almost universal support for the view that the positive EPP found in the literature are caused by the most productive firms self-selecting into export markets (for other recent studies see e.g., Kneller and Pisu, 2010; Vincenzo and Wagner, 2011). In the same spirit, Bernard et al. (2007, p. 111) argue that: “Results from virtually every study across industries and countries confirm that high productivity precedes entry into export markets”. At the same time, there is some empirical support for the hypothesis that

firms become more productive as a consequence of exporting (see e.g., De Loecker, 2007; Kneller and Pisu, 2010; Marin and Voigtländer, 2013). Of course the two hypotheses, self-selection and learning through exports, are not mutually exclusive. Accordingly, in a robustness analysis we employ a lag structure regressing lagged size and productivity on contemporaneous exporter status in order to focus – in line with the theory employed – on selection into exporting.

3.1 Data

Our data set consists of Danish firm-level data provided by Statistics Denmark for the period 1995-2012 and combines destination-specific export information (dating back to 1992) with business account information. Starting from the universe of all Danish firms, we exclude non-manufacturing firms and firms with missing sales or total sales below DKK 100,000 (about USD 17,000) per year. Furthermore, we require that firms have at least one employee. Thus, we only exclude extremely small if not erroneous firms but abstain from trimming the sample with respect to output, productivity or other firm characteristics. This allows us to maintain information on essentially all active manufacturing firms. In a given year a firm is classified as an exporter whenever merchandise exports are recorded and the declaration thresholds for extra- respectively intra-EU trade are exceeded.¹⁶ In the robustness analysis we apply an alternative definition: A firm is classified as an exporter if it exports for at least three successive periods. Accordingly, we have to exclude firms with less than three successive observation periods (between 1992 and 2012) from the analysis to prevent left truncation in the export status coding.

The resulting sample is composed of 19,767 firms, of which 7,758 at some point between 1995 and 2012 are classified as exporters. Capital and labour are measured as firms' total fixed assets and as firms' number of full-time equivalent employees, respectively. Descriptive statistics are provided in Table 1. Sectoral information is calculated at the 3-digit level, arriving at 89 sectors over 18 years with active exporters to compare and analyze sectoral EPP differences across.¹⁷

To briefly access how closely the productivity distribution for Denmark

¹⁶For Intra-EU exports we apply a common threshold of DKK 5.2 million for aggregate firm-level exports for each and every year. For Extra-EU exports we apply a common threshold of DKK 7,500 DKK per transaction.

¹⁷During the sample period the industry classification scheme changed from NACE 1.1 to NACE 2.0. We use overlapping information for both classification schemes for the years 2000-2008 to define firms' industry at the 3-digit NACE 2.0 level.

Table 1: **Descriptive Statistics**

	Mean	SD
ln Sales	15.93	1.61
ln Value Added	15.03	1.58
ln Capital	13.91	4.89
ln Labour	2.34	1.30
Exporter Dummy	0.41	0.49
<hr/>		
Observations	148824	
Firms	19767	

resembles a Pareto distribution, we transform Equation (3) into logs and obtain a convenient linear regression model that allows us to evaluate the goodness of fit of the Pareto distribution:

$$\ln(1 - \widehat{G}_j(\varphi_{ijt})) = \alpha - k_j \ln \varphi_{ijt} + \epsilon_{ijt} \quad (13)$$

with $\widehat{G}_j(\varphi_{ijt})$ denoting the empirical distribution function of productivity in industry j , i denoting firm and t time.

Figure 1 depicts plots of $\ln(1 - \widehat{G}(\varphi))$ against $\ln \varphi$ pooled over all industries, firms and years and the corresponding fitted lines and R^2 s as goodness of fit measures with φ being constructed as simple value added over workers.¹⁸ If productivity indeed follows a Pareto distribution, the depicted log-log relationship should follow a linear function. As found in previous literature, a linear function is not a good fit when considering all firms. However, when ordering firms according to their productivity and subsequently excluding the bottom 5th, 10th and 25th percentile, the linear fit between $\ln(1 - \widehat{G}(\varphi))$ and $\ln \varphi$ improves dramatically. Thus, the right tail of the actual productivity distribution has the best match to the Pareto.

3.2 Empirical Model

When taking theory to the data, it is important to acknowledge the wide range of forces that affect real countries and firms' trading patterns, but

¹⁸Note that in the theoretical model labour is the only production factor. When examining total factor productivity or domestic sales instead, the distribution turns out to be very similar.

Figure 1: **Productivity Distribution Pooled over all Industries**

that have been conveniently ignored in the theoretical set-up. For example, alternative drivers of trade, such as comparative advantages or effects stemming from other country asymmetries, are all absent in our theoretical considerations, but will matter in the data. Most importantly, firms will often not be single-product firms, as well as there might be firm-specific demand or quality differences. Furthermore, firm capital and purchased materials are relevant and unobserved firm idiosyncrasies may also affect firms' production technology and need to be accounted for. Finally, time and timing – an illusive concept when studying steady state equilibria as in Section 2 – will clearly be present in our firm-level data. While staying true to the theoretical model, our empirical analysis has to account for these realities.

Our baseline model reads:

$$\ln Z_{ijt} = \alpha_i + \delta \text{Exp}_{it} + \eta V_{jt} + \theta(\text{Exp}_{it} V_{jt}) + \beta_y \text{Year}_t + \epsilon_{ijt}, \quad (14)$$

where Z_{it} either is total sales of firm i at time t or its productivity measured as labour productivity (value added over full-time equivalent workforce) and total factor productivity (TFP), respectively. Our main variable of interest,

exporter status, is represented by the dummy variable Exp_{it} while V_{jt} represents industry j specific characteristics as outlined in Section 2, for example the degree of heterogeneity in productivity among the firms in the industry with $i \in j$. The parameter α_i refers to a firm-specific fixed effect controlling for firm idiosyncrasies that may determine our outcome of interest while being correlated with, e.g., exporter status or industry characteristics V_{jt} . Thus, this specification captures time-invariant firm-specific demand differences (or quality differences) through the firm fixed effect. Moreover, through the year dummies it accounts for common time-varying demand shocks, such as business cycle developments. Accordingly, in (14) the exporter premium is identified through longitudinal within-firm variation. The error term ϵ_{ijt} is allowed to be heteroscedastic and correlated within 3-digit industries j .

TFP is estimated applying the structural GMM estimator of Akerberg et al. (2015) drawing on intermediate goods purchases to invert unobserved firm-time-specific productivity. Moreover, we follow De Loecker (2007, 2013) and allow export status to affect subsequent firm productivity thereby avoiding bias of production function parameters. The corresponding estimations are carried out separately for each and every 2-digit industry to allow for industry-specific technologies.

Based on the parameter estimates obtained from estimating Equation 14, we quantify the relative size and productivity differences between exporters and non-exporters in per cent, i.e. we estimate the exporter premium $\widehat{EP}$ according to:

$$\widehat{EP}_{jt} = (\exp(\delta + \theta V_{jt}) - 1) * 100. \quad (15)$$

We obtain structural estimates of industry characteristics V_{jt} (which are not directly observable) by computing moments from the model in Section 2 and equating these moments to empirical observable counterparts in the data.¹⁹

The degree of heterogeneity in the productivity distribution is measured by the industries standard deviation of log labour productivity. This is a standard measure of productivity dispersion (see e.g. Bernard et al., 2003) and in our theoretical model this measure reads

$$Std_{jt}(\log \varphi | \varphi \geq \varphi_{jt,d}^*) = k_{jt}^{-1}. \quad (16)$$

Accordingly, the empirically observable productivity dispersion and our theoretical approach are comparable.

¹⁹Derivations of the moments presented below are elaborated in Appendix D.

In contrast to our measure of k_{jt} , directly observable empirical measures corresponding to σ_{jt} , $\frac{f_{jt,x}}{f_{jt}}$, and τ_{jt} are somewhat less straightforward to obtain.

For the elasticity of substitution σ_{jt} , we utilize that the coefficient of variation for sales of domestic firms on the domestic market (exporters as well as non-exporters) in the model from Section 2 reads

$$CV_{jt}(s_{jt,d}(\varphi) | \varphi > \varphi_{jt,d}^*) = \frac{\sigma_{jt} - 1}{\sqrt{k_{jt}(k_{jt} - 2(\sigma_{jt} - 1))}}. \quad (17)$$

Given the estimate of k_{jt} from above and the empirical coefficient of variation in the data, we obtain an estimate of σ_{jt} .

Authors such as Manova and Zhang (2012) and Di Comite et al. (2014) vividly emphasize the importance of idiosyncratic firm-product-destination specific demand shifters for explaining the variation in export prices and sales pointing to kaleidoscopic elasticities of substitution that vary between firms, products and destinations. To the extent that such idiosyncratic firm-product-destination-specific demand shifters are non-random, parameter estimates of Equation 14 may be biased. Ideally, our model would control for these potentially confounding idiosyncrasies, for instance through a more comprehensive set of fixed effects as in Di Comite et al. (2014). However, as total sales and productivity in Equation 14 are firm-specific and thus constant across products and export destinations, such an approach is not warranted by the variation in our data. An alternative approach is to use a more flexible demand function in our structural model to allow for market-specific elasticities of substitution that will directly affect selection into exporting of specific products into specific destinations.²⁰ Such firm-product-destination-specific elasticities, however, would need to be aggregated at least at the firm level to be included in Equation 14 which most likely would lead to simultaneity bias thereby rendering such an approach unfeasible.²¹ Accordingly, the reader should be aware of our identifying assumption that σ only systematically varies between industries and over time and cautiously interpret our findings.

Next, we address the ratio of fixed export costs to fixed costs in the domestic market. While we have a two-country theoretical model, Danish firms in the data export to several countries with different export barriers.

²⁰To obtain market-specific elasticities of substitution, one could employ a variant of Equation 17 substituting firm-product-destination-specific export sales for domestic sales.

²¹The only weighting factors to obtain the aggregate we can think of are the respective firm-product-destination export sales which are in the logic of our structural model determined through selection into exporting and thus dependent on firm productivity and size.

For our purpose, the important division of firms is between exporters and non-exporters, and by evoking a hierarchy of markets assumption, i.e. the assumption that firms enter the same export destinations in the same order, we can exploit the performance of Danish firms in the first destination, in order to estimate the trade barriers. The European Union plus Norway and Greenland (EU+) are the most obvious first markets for Danish exporters. Jointly these markets account for two thirds of all Danish exports. We thus aggregate the EU, Norway, and Greenland and classify them as a single primary export destination.²² We compute the ratio of the average sales in the domestic market across all firms, $\bar{s}_{jt,d}$, and the average sales in the primary export market (EU+), $\bar{s}_{jt,x}^{EU+}$ for the group of firms exporting to that market. This empirically observable ratio is in the model from Section 2 identical to the ratio of fixed export costs to fixed costs $\frac{f_{jx}}{f_j}$, i.e. the ratio we are interested in:

$$\frac{\bar{s}_{jt,x}^{EU+}}{\bar{s}_{jt,d}} = \frac{f_{jt,x}^{EU+}}{f_{jt}} \quad (18)$$

For the iceberg trade costs we exploit that the model predicts that the share of firms exporting to the primary market (cf. above), $S_{jt,x}^{EU+}$, reads

$$S_{jt,x}^{EU+} = \left(\frac{f_{jt,x}^{EU+}}{f_{jt}} \right)^{-\frac{k_{jt}}{\sigma_{jt}-1}} (\tau_{jt}^{EU+})^{-k_{jt}} \quad (19)$$

and given the estimates of $\frac{f_{jt,x}^{EU+}}{f_{jt}}$, σ_{jt} and k_{jt} the estimate of the iceberg trade costs τ_{jt}^{EU+} follows from the empirical share of firms exporting to the EU+. Table 2 provides summary statistics for the obtained deep parameters across our 89 sectors using equations (16) to (19).

4 Results

In what follows we focus on how ESP and EPP vary with the structural estimates of industry characteristics V obtained in Equations (16)–(19). We estimate variants of the baseline model stated in (14) allowing for firm-specific unobserved heterogeneity and the associated selection into exporter status and specific industries. To save space, Table 3 only reports $\widehat{EP}_{jt}$, i.e. the

²²Lawless (2009) shows that single-country destinations in the hierarchy of markets are only weakly supported in Irish data, hence we opt here for the aggregated EU+ as the primary destination.

Table 2: **Deep Parameters**

	Mean	SD	Bottom Decile	Median	Top Decile
k^{-1}	0.509	0.136	0.359	0.488	0.685
$\ln \tau$	0.448	1.100	-0.533	0.277	1.787
$\ln(f_x/f)$	0.204	0.954	-1.052	0.163	1.725
σ	1.998	0.332	1.693	1.976	2.360
Observations	148824				
Firms	19767				

outcome from (15), namely the estimated ESP and EPP.²³ According to Equation (15), ESP and EPP are allowed to differ along 3-digit industry characteristics V_{jt} . For this reason we estimate the premia at the mean, the bottom decile, the median, and the top decile of the respective estimates of industry characteristic V_j . We perform a Wald test on the H_0 that the premia for the top and bottom deciles are equal.

First, focussing on the estimate regarding industry-level productivity heterogeneity k^{-1} , we find a clear pattern in line with Corollary 1 that is statistically significant. In industries where firms are most homogeneous in terms of productivity, exporting firms on average have 27% higher sales than non-exporting firms. In contrast, in industries with the highest productivity heterogeneity exporter status is associated with 49% higher sales.²⁴ Moreover, exporting firms in the industries that are most heterogeneous (the top decile of k^{-1}) have on average substantially higher EPP than firms in industries that are very homogenous. The respective difference in EPP between industries in the top and bottom deciles of k^{-1} is about 19 percentage points for labour productivity and 15.5 percentage points for TFP.

²³Regression output tables for the underlying econometric estimations are available upon request.

²⁴Note that we do not curtail the universe of Danish firms but, different from e.g. ISGEP (2008), we even include very small unproductive and typically non-exporting firms, this lifts the general level of premia.

Table 3: ESP and EPP estimates, in per cent

Sales				
Mean	10th	50th	90th	Wald-Test
Corollary 1: $V \equiv$ productivity heterogeneity k_{jt}^{-1}				
36.53*** (2.21)	26.74*** (2.85)	35.12*** (2.27)	48.89*** (2.53)	47.71*** (p=0.00)
Corollary 2: $V \equiv$ ln iceberg costs τ_{jt}				
37.77*** (2.16)	41.84*** (2.63)	38.47*** (2.17)	32.40*** (2.90)	7.58*** (p=0.01)
Corollary 3: $V \equiv \ln \frac{f_{jt,x}}{f_{jt}}$				
38.93*** (2.19)	30.54*** (3.13)	38.65*** (2.17)	49.80*** (4.94)	8.07*** (p=0.00)
Corollary 4: $V \equiv \sigma_{jt}$				
36.55*** (2.42)	43.86*** (2.57)	37.06*** (2.38)	28.33*** (3.67)	14.84*** (p=0.00)
Observations 148824 Firms 19767				
Labour Productivity				
Mean	10th	50th	90th	Wald-Test
Corollary 1: $V \equiv$ productivity heterogeneity k_{jt}^{-1}				
7.09*** (0.74)	-1.16*** (1.10)	5.90*** (0.72)	17.58*** (1.70)	62.46*** (p=0.00)
Corollary 2: $V \equiv$ ln iceberg costs τ_{jt}				
7.97*** (0.89)	7.79*** (1.30)	7.94*** (0.93)	8.21*** (1.33)	0.05 (p=0.83)
Corollary 3: $V \equiv \ln \frac{f_{jt,x}}{f_{jt}}$				
8.16*** (0.88)	7.01*** (1.56)	8.12*** (0.86)	9.56*** (2.45)	0.49 (p=0.48)
Corollary 5: $V \equiv \sigma_{jt}$				
6.97*** (0.76)	14.41*** (1.56)	7.48*** (0.74)	-1.27 (1.94)	24.40*** (p=0.00)
Observations 148824 Firms 19767				
TFP, Akerberg et al. (2015)				
Mean	10th	50th	90th	Wald-Test
Corollary 1: $V \equiv$ productivity heterogeneity k_{jt}^{-1}				
3.84*** (0.70)	-3.02*** (0.92)	2.89*** (0.67)	12.48*** (1.52)	59.91*** (p=0.00)
Corollary 2: $V \equiv$ ln iceberg costs τ_{jt}				
4.52*** (0.90)	3.64*** (1.27)	4.37*** (0.94)	5.77*** (1.17)	1.57 (p=0.21)
Corollary 3: $V \equiv \ln \frac{f_{jt,x}}{f_{jt}}$				
4.52*** (0.86)	4.86*** (1.40)	4.53*** (0.84)	4.11*** (2.00)	0.06 (p=0.80)
Corollary 5: $V \equiv \sigma_{jt}$				
3.75*** (0.73)	9.76*** (1.46)	4.03*** (0.72)	-3.25 (1.81)	19.55*** (p=0.00)
Observations 105304 Firms 14546				

Notes: 10th, 50th, 90th are bottom, median, top decile. Standard errors in parentheses. *, **, *** statistically significant at the 10%, the 5%, the 1% levels, respectively. Firm-, time-fixed effects included. TFP estimation with structural GMM (Akerberg, Caves and Frazer (2006, 2015) and De Loecker (2007,2013) modification.

Second, we evaluate the relevance of variable trade costs τ for ESP and EPP. In terms of sales our findings are in contrast to Corollary 2 indicating that ESP are significantly lower for higher τ . Results for labour productivity and TFP point to a flat pattern. As indicated by the Wald tests reported in Table 3, EPP do not significantly vary with τ .

Third, we consider relative fixed export costs $\frac{f_x}{f}$. For the ESP, patterns follow Corollary 3. Industries in the bottom decile of $\frac{f_x}{f}$ on average are associated with lower ESP than industries with high relative fixed export costs, i.e. the top decile of $\frac{f_x}{f}$. As reported in Table 3, firms in industries in the bottom decile of $\frac{f_x}{f}$ on average show 31% higher sales than non-exporters whereas ESP increase to 50% for firms in the top decile industries. Differences in EPP between firms in industries in the top and bottom deciles of $\frac{f_x}{f}$ are much less pronounced and rendered statistically insignificant.

Fourth, we regard differentials between exporters and non-exporters subject to industry-specific structural estimates of the elasticity of demand σ . Corollary 4 had a clear prediction for ESP, while Corollary 5 – dealing with EPP – had an ambiguous prediction. Our empirical results strongly go against Corollary 4 as ESP are found to significantly fall in σ . As reported in Table 3, exporting firms in industries in the bottom decile of σ on average have 44% higher sales than non-exporters, an ESP that is dropping to 28% for firms in the top decile of σ . Moreover, while effects for the EPP theoretically are ambiguous, we find that in our data EPP unambiguously fall in σ . Labour productivity of exporting firms in the bottom decile of σ is 14% higher than that of non-exporting firms while EPP drop to 0 % in the top decile of σ . In terms of TFP the difference in EPP between the bottom and top deciles of σ is about 10 percentage points. These patterns are strongly statistically significant as indicated by the Wald-tests reported in Table 3.

To sum up: Based on an extension of the Melitz (2003) model we have been able to identify some underlying drivers of exporter premia size differences. Prominently, we find that ESP and EPP significantly rise in industry-level productivity heterogeneity k^{-1} . Furthermore, our empirical results challenge the positive relation between ESP and EPP and variable trade costs suggested by theory. If anything, ESP and EPP decrease in τ . Furthermore, there is some support for a positive relation between ESP and relative fixed export costs $\frac{f_x}{f}$. Finally, we find clear empirical support for a negative relation between ESP and EPP and the industry-specific elasticity of demand σ .

4.1 Robustness Tests

We estimate a number of alternative model specifications to assess the robustness of our findings. Table 4 only reports ESP and EPP estimates for the bottom and top deciles of V and the associated Wald-tests.

To start with, we estimate Equation 14 lagging the dependent variable by one period. This way we focus on the productivity-based selection into exporter status, akin to our theoretical model, as opposed to causal productivity effects of exporting since contemporaneous exporter status has no effect on past productivity. Furthermore, lagging sales and productivity by one period avoids potential simultaneity between the dependent variable and industry characteristics V . As shown in Specification (a) of Table 4, our earlier findings are not sensitive to these changes and identified ESP and EPP patterns remain qualitatively unchallenged.

We proceed by allowing for richer unobserved characteristics by estimating Equation 14 in first differences. Accordingly, we control for firm-specific unobserved trends in sales or productivity. Our approach amounts to a diff-in-diff-in-diff specification identifying ESP and EPP by comparing exporting to non-exporting firms that are on the same sales and productivity trajectory. Our results with regard to the relevance of k^{-1} , $\frac{f_x}{f}$ and σ in explaining ESP and EPP differences are qualitatively not altered as shown in specification (b) of Table 4. With respect to τ , however, we find no evidence for a reverse ESP pattern contrasting our main model specification. In terms of labour productivity there is even weak evidence suggesting a positive correlation between EPP and τ which would be in line with Corollary 2.

Finally, we alter our definition of exporter status reflecting on authors such as Békés and Muraközy (2012) who show that a large proportion of export spells are indeed only short-term with the least productive firms being most likely to stop exporting. Our alternative exporter definition now requires that a firm exports for at least three successive periods. As reported in specification (c) of Table 4, this alternative exporter definition does not qualitatively change our earlier findings. Again, there is weak evidence that EPP increase with τ in line with Corollary 2. However, this only holds for one of the two productivity measures.

Table 4: Robustness Tests, ESP and EPP estimates in per cent

Sales								
(a) lag structure			(b) firm-specific trends			(c) long-term exporter		
10th	90th	Wald-Test	10th	90th	Wald-Test	10th	90th	Wald-Test
Corollary 1: $V \equiv$ productivity heterogeneity k_{jt}^{-1}								
18.03*** (2.23)	31.97*** (1.97)	26.49*** (p=0.00)	5.02*** (0.72)	8.81*** (0.83)	19.08*** (p=0.00)	39.23*** (3.63)	61.72*** (3.36)	41.02*** (p=0.00)
Corollary 2: $V \equiv$ Log iceberg costs τ_{jt}								
28.47*** (2.06)	20.42*** (1.66)	14.60*** (p=0.00)	7.06*** (0.85)	6.58*** (0.83)	0.20 (p=0.65)	54.96*** (3.46)	44.41*** (3.76)	6.76** (p=0.01)
Corollary 3: $V \equiv \text{Log } \frac{f_{jt,x}}{f_{jt}}$								
19.23*** (1.93)	34.32*** (3.51)	11.95*** (p=0.00)	6.67*** (0.97)	7.11*** (1.26)	0.06 (p=0.81)	41.58*** (3.98)	65.47*** (6.31)	8.24*** (p=0.00)
Corollary 4: $V \equiv \sigma_{jt}$								
28.89*** (1.82)	18.91*** (2.66)	12.69*** (p=0.00)	7.89*** (0.79)	5.36*** (0.86)	6.30*** (p=0.01)	56.57*** (3.28)	40.95*** (4.41)	13.90*** (p=0.00)
Observations 127554			Observations 127554			Observations 148824		
Labour Productivity								
(a) lag structure			(b) firm-specific trends			(c) long-term exporter		
10th	90th	Wald-Test	10th	90th	Wald-Test	10th	90th	Wald-Test
Corollary 1: $V \equiv$ productivity heterogeneity k_{jt}^{-1}								
2.21 *** (1.07)	8.25 *** (0.92)	15.37 *** (p= 0.00)	-2.01 *** (0.77)	5.88 *** (1.18)	25.12 *** (p= 0.00)	-0.09 (1.26)	18.75 *** (1.86)	62.68 *** (p= 0.00)
Corollary 2: $V \equiv$ Log iceberg costs τ_{jt}								
5.96 *** (0.93)	4.25 *** (0.81)	1.78 (p= 0.18)	0.70 (0.97)	3.06 *** (0.83)	3.46 * (p= 0.06)	8.66 *** (1.45)	9.82 *** (1.64)	0.30 (p= 0.58)
Corollary 3: $V \equiv \text{Log } \frac{f_{jt,x}}{f_{jt}}$								
3.91 *** (1.01)	7.27 *** (1.78)	1.80 (p= 0.18)	2.01 *** (0.79)	1.31 (1.20)	0.20 (p= 0.65)	8.64 *** (1.96)	10.04 *** (2.83)	0.11 (p= 0.74)
Corollary 5: $V \equiv \sigma_{jt}$								
7.01 *** (0.72)	2.50 *** (1.16)	11.01 *** (p= 0.00)	4.75 *** (1.05)	-2.40 *** (1.12)	14.97 *** (p= 0.00)	15.56 *** (1.65)	-0.17 (2.01)	26.00 *** (p= 0.00)
Observations 127554			Observations 127554			Observations 148824		
TFP, Akerberg et al. (2015)								
(a) lag structure			(b) firm-specific trends			(c) long-term exporter		
10th	90th	Wald-Test	10th	90th	Wald-Test	10th	90th	Wald-Test
Corollary 1: $V \equiv$ productivity heterogeneity k_{jt}^{-1}								
0.00 (1.02)	4.29*** (1.01)	8.56*** (p=0.00)	-1.81** (0.75)	4.65*** (1.12)	19.54*** (p=0.00)	-3.01*** (1.21)	12.14*** (1.68)	57.08*** (p=0.00)
Corollary 2: $V \equiv$ Log iceberg costs τ_{jt}								
2.27** (1.02)	1.93*** (0.85)	0.07 (p=0.79)	0.34 (1.05)	2.43*** (0.83)	2.23 (p=0.14)	3.20*** (1.45)	6.35*** (1.56)	3.12* (p=0.08)
Corollary 3: $V \equiv \text{Log } \frac{f_{jt,x}}{f_{jt}}$								
1.33 (1.03)	3.35* (1.81)	0.65 (p=0.42)	1.33 (0.83)	1.14* (1.42)	0.01 (p=0.92)	5.29*** (1.88)	3.35 (2.21)	0.32 (p=0.57)
Corollary 5: $V \equiv \sigma_{jt}$								
3.87*** (0.88)	-0.57 (1.16)	9.23*** (p=0.00)	3.63*** (1.01)	-2.14* (1.09)	10.79*** (p=0.00)	9.56*** (1.57)	-3.30* (1.92)	21.04*** (p=0.00)
Observations 108589			Observations 108589			Observations 127000		

Notes: 10th, 50th, 90th are bottom, median, top decile. Standard errors in parentheses. *, **, *** statistically significant at the 10%, the 5%, the 1% levels, respectively. Firm-, time-fixed effects included. TFP estimation with structural GMM (Akerberg, Caves and Frazer (2006, 2015) and De Loecker (2007,2013) modification.

5 Discussion

It is worth to explicitly state that it is not the aim of this paper to test the Melitz (2003) model or extension hereof, nor is it the paper’s aim to show how to avoid bias when estimating ESP or EPP. Industry-specific exporter premia can be estimated without conditioning on industry-specific characteristics. Including controls for all industry-specific characteristics (e.g., by including industry-time-specific constants) would result in an over-specified model understating true exporter premia that according to our theory exactly depend on some of these characteristics. However, the important issue is that our results show that when it comes to solving the puzzle of why exporter premia differ so widely across industries, one has to take into account the various underlying industry characteristics, as illustrated by our theoretical framework. As such, our paper has identified four relevant parameters that matter for the magnitude of observed premia: industry-level productivity heterogeneity, variable trade costs, relative fixed costs of exporting and the sectoral elasticity of demand. Obviously, this in turn implies that the origins for sector differences in these characteristics are important future alleys for empirical research.

In order to demonstrate how much of the cross-industry variation in exporter premia in the data can be explained by the different industry characteristics that we have identified, we carry out a simple purely descriptive exercise. Instead of explicitly modeling the interaction of ESP and EPP with our four deep parameters, we, in a first step, obtain average sectoral-level $\overline{ESP}_j$ and $\overline{EPP}_j$ (one for each j of our 3-digit industries). We do this by estimating the model in (14) sector by sector, but without interaction terms. Our sectoral-level ESP and EPP are thus conditional on firm-level unobserved heterogeneity, year- and industry-specific shocks.²⁵

In a second step we regress sectoral average $\overline{ESP}_j$ and $\overline{EPP}_j$ on the 1995 values of the industry-level deep parameters $V_{j,1995}$. The corresponding R^2 s will reflect how much of the total variation in ESP and EPP can be explained by the initial values of the deep parameters suggested by theory. The resulting R^2 s are reported in Table 5. This approach is merely a back of the envelope calculation as, different from the models estimated in Section 3, it does not correct for the systematic selection of heterogeneous firms into industries with specific characteristics. However, from the exercise we may cautiously conclude that about 13% of the cross-industry variation in ESP stem from cross-industry differences in the four deep parameters identified

²⁵To estimate industry-by-industry ESP and EPP, we aggregate 3-digit industries 267 and 268 (NACE 2.0) to obtain sufficient industry-level observations.

in the theory. For EPP in terms of labour productivity, the respective figure is 26%. For our most comprehensive EPP measure based on the structural GMM estimator of Akerberg et al. (2015) that accounts for firm- and time-specific productivity shocks, our industry-level deep parameters can even explain 43 per cent of the cross-industry variation in EPP. Thus, we conclude that a sizable share of the observed variation of ESP and EPP across industries is explained by cross-industry differences of four profound industry characteristics.

Table 5: **Explaining ESP and EPP differences**

	R^2		
	$\overline{ESP}_j$ Sales	$\overline{EPP}_j$ Labour Productivity	$\overline{EPP}_j$ Akerberg et al.(2015)
$k_{j,t=1995}^{-1}, \ln \tau_{j,t=1995}$ $\ln(f_x/f)_{j,t=1995}, \sigma_{j,t=1995}$	0.13	0.26	0.43
Observations	89	89	89

It is remarkable that the present approach can capture a sizable share of the observed variation. In particular, because the structural estimation approach applied here to measure industry characteristics is clearly not the only conceivable way to link our model insights to empirical data. An alternative approach would be to harvest other proxies that resemble the deep parameters identified in our framework. For example, instead of backing out the estimates for trade costs via the structural estimation, standard gravity equation type measures could be applied, say a trade weighted average measure of export costs across all destinations for the various sectors. Still, since international trade in itself is a function of export costs, this alternative route has a range of problems that would have to be addressed.

Importantly, the above results are derived relying on a specific version of the Melitz (2003) model. This has the advantage of embedding our empirical investigation in the mainstream of international trade theories with heterogenous firm exports. However, the approach also has costs. Staying true to the model ignores a range of factors that recent empirical work has uncovered. For example the prevalence of multiproduct firms and non-constant mark-ups. Yet, the literature has made substantial advances in capturing these aspects both formally and empirically, see for example Eckel and Neary (2010), Eckel et al. (2016). Multiproduct firms complicate matters. The reason being that most firm-level variables are weighted averages of

firm-product-level variables (e.g. productivity and sales) while the firm-level export indicator takes the value one even if only a single product from the product portfolio of the firm is exported. Changes in the environment of the firm (e.g. the deep parameters identified in this paper) may not only affect export status and firm-product level variables but also the relative weights of firm-product variables as well as the range of the firm-level product portfolio.

Other important empirical features not reflected in our model framework are firm-specific differences in demand and quality. Works such as Manova and Zhang (2012) and Di Comite et al. (2014) have established that quality and demand (taste) shifts play an important role in explaining for example firm sales. It is thus instructive to examine the robustness of our model framework, once it is extended to include firm-destination-specific demand shifters and firm-quality heterogeneity on top of heterogeneity in marginal production efficiency (φ). We provide such an extension in Appendix B and show that a limited set of additional assumptions suffices to re-establish the model predictions of the simpler framework from Section 2 even in the extended model.

6 Conclusion

Exporter premia (in productivity or size) are a central stylized fact of international economics, albeit the accumulated empirical evidence discloses substantial – and so far unexplained – EPP and ESP size differences. The present paper tackles this issue. First, we ask if at all – and in what direction – a simple workhorse model of heterogeneous firms trade (i.e. Melitz, 2003) contains predictions on the determinants of the magnitude of exporter premia. In particular, we introduce explicit EPP and ESP measures into a version of the Melitz (2003) model. This model gives a number of new insights and predictions. Second, we use this model to investigate variations in cross-sectoral exporter premia in Danish firm-level data. We are able to identify an important role for the degree of heterogeneity in productivity in explaining the variance in EPP and ESP magnitude. Both in the theoretical framework and empirically, industries where the realized productivity distribution displays a wider variance feature larger EPP and ESP. Similarly, our results link larger relative fixed export costs to larger exporter premia. Finally, in the model the sectoral elasticity of demand has ambiguous effects on the magnitude of the EPP and a positive correlation with the ESP. Yet, in the data we find strong empirical patterns that sectors with higher elasticity of demand display smaller EPP and smaller ESP.

Our results contribute several directions for future research. In terms of

theoretical extensions, for example, the relation between the elasticity of demand and the EPP and ESP found in our data should ideally, if confirmed by other empirical studies, feature in the models of international trade. Similarly, the link between fixed export costs (which matter in the policy debate surrounding trade facilitation) and the EPP and ESP should be further investigated. In particular, the prevalence of sizable exporter productivity and sales premia has sparked renewed interest in policies that assist the international activity of firms. In terms of the overall production efficiency in a sector or the economy, a larger exporter productivity premium makes it, everything else equal, more attractive to aim at industrial policies that reallocate production resources (market shares) from non-exporters towards incumbent exporters. Thus, optimal industrial policies (e.g. export guarantees vs. general production subsidies) may well depend on the magnitude of exporter productivity premia. Our paper has shown how the size of these premia in turn depends on a limited number of deep parameters. Moreover, our paper shows that certain policies, say export promotion programmes that constitute reductions in fixed export costs, will at least in the model we provided reduce exporter premia. Exploring the links between optimal policy design and industry observables, such as exporter premia, is an important topic for future research.

Most importantly, our findings provide new perspectives on the sizable empirical literature on exporter premia of the past decades. We highlight briefly two directions in which the present paper contributes. Firstly, we are able to identify central fundamentals that matter for size differences in exporter premia. We establish that the variance in the underlying productivity distribution, the relative size of fixed export costs and the sectoral elasticity of demand are decisive fundamentals for EPP and ESP realizations. These findings, if supported by other empirical studies, are worthy candidates for future stylized facts. Secondly, going beyond the narrow focus of sectoral variation considered in the present paper, our analysis suggests that the cross-country variation in exporter premia that is well established in the literature can at least partly be explained by measurable cross-country differences in industry structure and industry characteristics.

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