



Beyond Constant Markups: Identifying the Export Elasticity under Strategic Pricing

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Abstract

Standard estimators of the export price elasticity assume firms set prices independently of competitors, which yields inconsistent estimates under strategic interaction. We propose a heteroskedasticity-based estimator that relaxes this restriction by applying the Frisch-Waugh-Lovell theorem to partial out rival price and quantity aggregates, jointly identifying the export price elasticity and the strategic complementarity elasticity within a GMM estimation framework. The estimator requires no quasi-experimental variation and no data beyond standard bilateral trade flows and exchange rates. Applying the method to Norwegian manufacturing exports over 1999–2024, we estimate an export price elasticity of 4.42 (standard error of 0.72) and a strategic complementarity elasticity of 0.51 (standard error of 0.15). The significant positive complementarity estimate rejects the constant-markup restriction imposed by standard estimators.

Keywords: Export Elasticity, Panel Data, Two-Stage Estimator, GMM

JEL Classification: C13, C36, F14, L13

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Sammendrag

Standardestimater av eksportpriselastisiteten bygger på en antakelse om at bedrifter fastsetter priser uavhengig av konkurrentene. Dersom bedriftene derimot bruker strategisk prissetting vil denne antakelsen gi inkonsistente estimater for eksportpriselastisiteten. Vi utvikler en estimator basert på heteroskedastisitet som omgår denne antakelsen ved å anvende Frisch-Waugh-Lovell-teoremet til å kontrollere for konkurrentenes priser og kvanta. Dette muliggjør simultan identifikasjon av både eksportpriselastisiteten og elastisiteten for strategisk komplementaritet. Estimatoren krever verken kvasi-eksperimentell variasjon eller data utover bilateral handelsdata og valutakurser. Ved å bruke metoden på norsk industrieksport i perioden 1999–2024 estimerer vi en eksportpriselastisitet på 4,42 (standardfeil på 0,72) og en elastisitet for strategisk komplementaritet på 0,51 (standardfeil på 0,15). Det positive og statistisk signifikante estimatet på strategisk komplementaritet innebærer at antakelsen om konstant prispåslag (*markup*) kan forkastes.

1 Introduction

The export price elasticity governs how bilateral exports respond to trade costs and world prices, fundamentally shaping the welfare effects of trade policy (Hillberry and Hummels, 2013). Yet standard estimators resolve the simultaneity problem under a constant-markup restriction that is violated when firms price strategically. In this paper, we develop a heteroskedasticity-based estimator to accommodate strategic pricing and endogenous markups, jointly identifying the export price elasticity and the strategic complementarity elasticity without requiring external instruments.

Heteroskedasticity-based GMM estimators in the tradition of Feenstra (1994) typically combine CES demand with a reduced-form isoelastic supply relation, implicitly assuming constant markups (Feenstra, 1994; Imbs and Mejean, 2015; Grant and Soderbery, 2024; Brasch et al., 2025), a restriction maintained even in extensions that enrich the supply side with firm heterogeneity and increasing marginal market-penetration costs (Bergstrand et al., 2023). However, a vast literature documents strategic pricing and endogenous markup variation across markets and competitive environments (Krugman, 1987; Atkeson and Burstein, 2008; Burstein and Gopinath, 2014; Amiti et al., 2019; Burstein et al., 2025; Gagliardone et al., 2025). If markups vary endogenously with demand or supply conditions, the key identifying assumption of orthogonality between structural demand shocks and supply-side shocks is violated.

We restore identification by developing a structural model with CES demand and profit-maximizing exporters that set prices strategically, expanding the standard econometric system to include rival price and quantity aggregates. Strategic pricing distorts the identifying orthogonality condition through observable rival aggregates. However, partialling them out of the GMM estimation equation via the Frisch-Waugh-Lovell theorem restores identification without a need for external instruments. Our estimator requires neither quasi-experimental variation nor data beyond standard bilateral trade flows and exchange rates, and it identifies both the export price elasticity and the strategic complementarity elasticity. To our knowledge, this is the first application of heteroskedasticity-based identification to strategic price setting.

We apply the estimator to Norwegian manufacturing exports from 1999–2024, combining transaction-level customs data with bilateral trade data from CEPII-BACI. The estimation panel covers three-digit SITC product groups, yielding 6,007 product-year observations. We estimate an export price elasticity of 4.42 (standard error of 0.72) and a strategic complementarity elasticity of 0.51 (standard error of 0.15), both significant at the 1 percent level. The precisely estimated strategic complementarity elasticity rejects the constant-markup restriction underlying standard heteroskedasticity-based estimators.

The remainder of the paper is organized as follows: Section 2 outlines the CES pricing model at the destination level, describes aggregation across destinations, and summarizes the key parameters to be estimated. Section 3 constructs the estimator and states the identifying assumptions. Section 4 describes the data and presents the estimation results, and Section 5 concludes.

2 Model Framework: Demand, Pricing, and Aggregation

This section presents the model linking exporter-level prices and quantities to destination-specific demand conditions and pricing. We start from CES demand within each destination, then introduce a pricing block in which exporters set destination prices as a markup over marginal cost with markups responding to competitors' prices. We then aggregate destination-level objects to the exporter level, yielding an export-level demand curve and an export-level price-setting equation. The resulting system centers on two key param-

ters: the export price elasticity, σ , and the strategic complementarity elasticity, γ .

2.1 Environment

The economy consists of a finite set of exporters, $n \in \Sigma_{\text{exp}}$, each producing a single horizontally differentiated variety, and a set of destination markets, $d \in \Sigma_{\text{dest}}$. Exporters internalize their effect on the destination-specific industry price index, thereby giving rise to strategic price setting and variable markups. The index n refers to exporters, d to destination markets, t to time and g is a product index. In the theoretical sections, the product is fixed and the index g is suppressed, whereas in the empirical application (which focuses on Norway), n is fixed and g varies.

2.2 Destination-Specific Demand

The price of a good produced by exporter n and sold in destination d at time t is denoted by P_{dnt} , with corresponding quantity X_{dnt} . Aggregate destination-level quantity, X_{dt} , and price, P_{dt} , are CES-indices over all origins defined as follows:

$$P_{dt} = \left(\sum_n \alpha_{dn} P_{dnt}^{1-\sigma} \right)^{\frac{1}{1-\sigma}}, \quad (1)$$

$$X_{dt} = \left(\sum_n \alpha_{dn}^{\frac{1}{\sigma}} X_{dnt}^{\frac{\sigma-1}{\sigma}} \right)^{\frac{\sigma}{\sigma-1}}, \quad (2)$$

where $\sigma > 1$ is the Armington elasticity, i.e. the elasticity of substitution across origin varieties, n . The implicit assumption is that σ does not depend on destination, d . It measures the ease with which buyers in destination d substitute away from an origin variety whose delivered price rises toward relatively cheaper varieties. The parameters $\alpha_{dn} > 0$ are preference weights that sum to unity.

From the perspective of exporter n , the Armington elasticity σ corresponds directly to the price elasticity of export demand.¹ The demand for a traded variety originating from exporter n and consumed in destination d follows from Shephard's lemma:

$$\Delta \ln X_{dnt} = -\sigma \Delta \ln P_{dnt} + \sigma \Delta \ln P_{dt} + \Delta \ln X_{dt}. \quad (3)$$

This characterizes how quantity demanded responds to changes in own price and destination-level price and volume indices, reflecting standard CES properties under homothetic preferences.

To isolate exporter n 's own price from rival prices, we decompose the destination aggregates P_{dt} and X_{dt} into exporter n 's contribution and a leave-one-out foreign composite. Let $P_{dt}^{F(n)}$ denote the foreign price index in destination d and $X_{dt}^{F(n)}$ the corresponding volume index, both aggregated over exporter n 's competitors as Sato-Vartia log-change indices (Sato, 1976; Vartia, 1976):

$$\Delta \ln P_{dt}^{F(n)} = \sum_{m \neq n} \psi_{dmt}^{F(n)} \Delta \ln P_{dmt}, \quad \Delta \ln X_{dt}^{F(n)} = \sum_{m \neq n} \psi_{dmt}^{F(n)} \Delta \ln X_{dmt},$$

where $\psi_{dmt}^{F(n)}$ are log-mean expenditure shares of competitors $m \neq n$ in destination d , defined analytically in Section 4.2. The superscript $F(n)$ indicates that both the set of competitors and their trade value shares in

¹Under oligopolistic competition, this correspondence breaks down as demand elasticities vary with market shares (Atkeson and Burstein, 2008).

destination d depend on which exporter n is excluded from the aggregate. The destination aggregate indices are then aggregated from the exporter's contribution and the leave-one-out composite:

$$\Delta \ln P_{dt} = \omega_{dnt} \Delta \ln P_{dnt} + (1 - \omega_{dnt}) \Delta \ln P_{dt}^{F(n)}, \quad (4)$$

$$\Delta \ln X_{dt} = \omega_{dnt} \Delta \ln X_{dnt} + (1 - \omega_{dnt}) \Delta \ln X_{dt}^{F(n)}, \quad (5)$$

where ω_{dnt} is exporter n 's market share in destination d .² Substituting these decompositions into Equation (3) preserves its functional form:

$$\Delta \ln X_{dnt} = -\sigma \Delta \ln P_{dnt} + \sigma \Delta \ln P_{dt}^{F(n)} + \Delta \ln X_{dt}^{F(n)}. \quad (6)$$

This demand specification expresses quantity changes as a function of exporter n 's own price, foreign competitor prices, and foreign volume, isolating the competitive effects within each destination market.

2.3 Destination-Specific Price Setting

Exporters set prices strategically in response to competitive conditions. Each exporter n supplies its product to destination markets d , facing conditions that differ across destinations and shape the strategic environment for price setting. The accounting identity for log prices is:

$$\ln P_{dnt} = \ln M_{dnt} + \ln MC_{nt}, \quad (7)$$

which defines the realized markup (M) as the log difference between the firm's price and its marginal cost (MC). The markup is specified as a function of competitors' pricing behavior:

$$\ln M_{dnt} = m_{nt} + \Gamma \ln(P_{dt}^{F(n)} / P_{dnt}). \quad (8)$$

Markups adjust with rival prices at each destination because a change by one competitor shifts customer spending across products, altering demand sensitivities and thus the profit-maximizing margin for the exporter. Equation (8) characterizes the exporter's optimal markup conditional on the rival-price index; it does not imply that the exporter substantially affects that index. [Burstein and Gopinath \(2014\)](#) describe a number of models that produce this reduced-form relationship between markups and relative price. [Amiti et al. \(2019\)](#) show that the representation holds both on and off equilibrium: when each firm's price satisfies its corresponding fixed point given the prevailing prices of its competitors, the industry is in equilibrium, while off equilibrium, the same condition describes the firm's reaction function. Although our model is based on CES preferences, [Amiti et al. \(2019\)](#) also demonstrate that the framework accommodates both homothetic and non-homothetic preferences and extends to a broad class of oligopolistic environments including simultaneous-move Bertrand and Cournot competition as well as certain sequential-move settings such as Stackelberg competition.

To close the model on the cost side, we specify a marginal cost function that depends on aggregate exporter output:³

$$\ln MC_{nt} = K_{nt} + \Phi \ln X_{nt}, \quad K_{nt} > 0. \quad (9)$$

²Consistency in aggregation holds for the Montgomery-Vartia (Vartia I) index ([Vartia, 1976](#)) but not for the Sato-Vartia (Vartia II) index used here. The Sato-Vartia index is exact under CES aggregation and satisfies proportionality, time reversal, and the factor reversal test; the Montgomery-Vartia index fails proportionality. The empirical sensitivity to choice of index is assessed in [Appendix B](#).

³This follows the bilateral export supply specification in [Bergstrand et al. \(2023\)](#), adapted to aggregate rather than destination-specific output in the spirit of [Vannoorenberghe \(2012\)](#); see also [Arkolakis \(2010\)](#) for the underlying market-penetration cost microfoundation.

Because production takes place domestically, marginal costs depend on the scale of domestic production and are independent of the destination to which output is sold. The parameter K_{nt} captures baseline unit costs, including input prices and production efficiency. The elasticity Φ governs the response of marginal cost to output scale, reflecting the degree of returns to scale in production. The case $\Phi = 0$ corresponds to constant returns to scale. Positive Φ reflects diminishing returns to scale, while negative Φ reflects increasing returns to scale.

Substituting Equations (8) and (9) into Equation (7) yields the optimal destination price:

$$\ln P_{dnt} = \eta_{nt} + \phi \ln X_{nt} + \gamma \ln P_{dt}^{F(n)}, \quad (10)$$

where all composite parameters are attenuated by strategic interaction through the factor $(1 + \Gamma)^{-1}$:

$$\eta_{nt} = \frac{K_{nt} + m_{nt}}{1 + \Gamma}, \quad \phi = \frac{\Phi}{1 + \Gamma}, \quad \gamma = \frac{\Gamma}{1 + \Gamma}.$$

The exporter-time intercept η_{nt} combines baseline marginal costs, K_{nt} , and the baseline markup component, m_{nt} . The remaining destination-specific variation in prices arises through the rival-price index $P_{dt}^{F(n)}$. The composite strategic complementarity elasticity (γ) measures the pass-through from competitor price changes to the exporter's price. This widely used elasticity quantifies the strength of strategic complementarities in price setting and is typically positive in empirical applications.⁴ When $\gamma = 0$, the firm's markup is constant and prices are determined solely by marginal cost, as in standard monopolistic competition.

2.4 Aggregation Across Destinations

We now outline the framework for aggregating destination-specific export demand and price setting. Market-specific regulatory requirements, logistics, distribution networks, and contracting arrangements differentiate otherwise identical goods across destinations. Reallocating sales across destinations therefore changes marginal profitability, making destination-specific shipments imperfect substitutes from the exporter's perspective.⁵ We capture this imperfect substitutability by modeling the exporter's aggregate export quantity as a CES composite index over destination-specific shipments, with a constant elasticity of substitution across markets. This approach is closely related to the CES-based aggregation and exact-index foundations in Redding and Weinstein (2024), but it shifts the object of aggregation from firms/products (importer-side) to destination markets (exporter-side). Specifically, for each exporter n , aggregate export quantity is defined as

$$X_{nt} = \left(\sum_{d \in \Sigma_{\text{dest}}} \xi_{dn}^{\frac{1}{\rho_n}} X_{dnt}^{\frac{\rho_n - 1}{\rho_n}} \right)^{\frac{\rho_n}{\rho_n - 1}}, \quad \rho_n > 1, \quad (11)$$

where ξ_{dn} denote normalized destination-specific preference or quality weights that sum to one. The CES aggregator nests standard linear aggregation in the limit as the elasticity of substitution approaches infinity and quality weights become uniform across destinations. The corresponding CES price index consistent with this aggregation is

$$P_{nt} = \left(\sum_{d \in \Sigma_{\text{dest}}} \xi_{dn} P_{dnt}^{1 - \rho_n} \right)^{\frac{1}{1 - \rho_n}}. \quad (12)$$

⁴ Amiti et al. (2019) note that γ can be negative in principle, implying strategic substitutes where firms lower prices in response to competitor price increases.

⁵ Rauch (1999) establishes that market-specific characteristics create effective product differentiation across bilateral trading relationships, even for physically identical goods.

Under the CES aggregation structure, the exact discrete-time changes in aggregate quantity and the corresponding price index admit a Sato-Vartia representation. Hence,

$$\Delta \ln P_{nt} = \sum_{d \in \Sigma_{\text{dest}}} \psi_{dnt} \Delta \ln P_{dnt}, \quad (13)$$

$$\Delta \ln X_{nt} = \sum_{d \in \Sigma_{\text{dest}}} \psi_{dnt} \Delta \ln X_{dnt}, \quad (14)$$

where the weights ψ_{dnt} are exporter n 's destination shares.⁶ The Sato-Vartia representation is exact under CES aggregation and superlative, providing a locally robust approximation for modest departures from the CES benchmark (Barnett and Choi, 2008).⁷

Next, we construct exporter- n -specific rival price and quantity aggregates as

$$\Delta \ln P_{nt}^F = \sum_{d \in \Sigma_{\text{dest}}} \psi_{dnt} \Delta \ln P_{dt}^{F(n)}, \quad (15)$$

$$\Delta \ln X_{nt}^F = \sum_{d \in \Sigma_{\text{dest}}} \psi_{dnt} \Delta \ln X_{dt}^{F(n)}. \quad (16)$$

These aggregates vary across exporters because common destination-level conditions are weighted by exporter-specific destination shares. An exporter concentrated in a few markets is therefore more exposed to price and quantity movements in those markets than a more diversified exporter.

2.5 Export-Level Demand and Price Setting Curves

Aggregating across destinations using the destination-specific demand function in Equation (6), the aggregation indices in Equations (13) and (14), together with the definitions of the competitive price and demand indices in Equation (15) and Equation (16), yields the exporter-level demand curve for a given exporter:

$$\Delta \ln X_{nt} = -\sigma \Delta \ln P_{nt} + \sigma \Delta \ln P_{nt}^F + \Delta \ln X_{nt}^F. \quad (17)$$

The export demand curve entails constant proportional elasticities of export quantities with respect to both own and foreign prices and homothetic scaling with foreign quantity. Because the Sato-Vartia indices are weighted sums of destination-level log changes and σ is common across destinations, σ factors out of the aggregation. The exporter-level demand curve therefore retains the destination-level CES form as in Equation (3).

The exporter-level price setting curve is obtained by first-differencing the optimal log-linear price equation in Equation (10) and aggregating destination prices across markets using the Sato-Vartia price index in Equation (13). Specifically, the export-level price-setting equation takes the form

$$\Delta \ln P_{nt} = \Delta \eta_{nt} + \phi \Delta \ln X_{nt} + \gamma \Delta \ln P_{nt}^F. \quad (18)$$

This curve captures how an exporter's aggregate price responds to marginal costs, which vary with the scale of production, and rival prices.

Equations (17) and (18) form the core empirical system. The first governs how export quantities respond to relative prices and foreign market conditions; the second governs how exporters adjust prices in response to costs and competitor pricing.

⁶We define the Sato-Vartia weights analytically in the empirical application in Section 4.

⁷Under linear aggregation, which arises as the limiting case of the CES aggregator as $\rho_n \rightarrow \infty$, the weights ψ_{dnt} reduce to log-mean quantity shares, so each destination's contribution to aggregate export growth is proportional to its quantity share. For export price measurement, however, unit values may be misleading when cross-destination variation reflects differences in product characteristics or terms of sale rather than pure price changes (Silver, 2010).

3 Identification and Empirical Implementation

The simultaneous determination of export demand and pricing in Equations (17) and (18) implies endogeneity problems when estimating both σ and γ . Identifying γ from the price-setting curve requires an instrument for marginal costs that is orthogonal to demand shocks.⁸ Identifying σ from the demand curve presents a distinct problem: observed quantity changes conflate the causal effect of price on demand with the feedback from output to marginal costs and prices. Standard heteroskedasticity-based estimators resolve both problems under the restriction $\gamma = 0$, but are inconsistent when strategic pricing is present. We relax this restriction by augmenting the moment conditions with exporter-level rival price and quantity aggregates, as formalized below.

3.1 Heteroskedasticity-Based Identification: The Baseline Framework

Heteroskedasticity-based estimators in the tradition of [Feenstra \(1994\)](#) identify structural elasticities by exploiting cross-sectional variation in supply and demand shocks, thereby eliminating the need for external instruments. The canonical framework specifies demand and price determination in a simultaneous two-equation system and implicitly assumes constant markups.⁹ Given parameters $\beta = 1 - \sigma$ and α , the baseline specification is:

$$\begin{aligned}\ddot{\Delta} \ln V_{nt} &= \beta \ddot{\Delta} \ln P_{nt} + \ddot{\Delta} \varepsilon_{nt}^D, \\ \ddot{\Delta} \ln P_{nt} &= \alpha \ddot{\Delta} \ln V_{nt} + \ddot{\Delta} \varepsilon_{nt}^S,\end{aligned}\tag{19}$$

where V_{nt} denotes export value, P_{nt} denotes the export price, ε_{nt}^D and ε_{nt}^S are structural demand and supply disturbances, and the operator $\ddot{\Delta}$ removes additive exporter and time fixed effects via two-way differencing ([Wooldridge, 2025](#)): for any variable z_{nt} ,

$$\ddot{\Delta} z_{nt} = \Delta z_{nt} - \Delta \bar{z}_{.t},$$

where $\Delta z_{nt} = z_{nt} - z_{n,t-1}$ is the ordinary first difference and $\bar{z}_{.t} = N^{-1} \sum_{n=1}^N z_{nt}$ is the cross-sectional mean.¹⁰ This simultaneous two-equation system, identified by exploiting cross-sectional heteroskedasticity in the structural shocks, underlies standard heteroskedasticity-based estimators in the literature.

Identification rests on the assumption that demand and supply disturbances are mean-zero and mutually orthogonal, i.e. $\mathbb{E}[\varepsilon_{nt}^D \varepsilon_{nt}^S] = 0$ for all n and t . This requires that demand shocks, such as idiosyncratic shifts in destination preferences, are uncorrelated with supply shocks, such as exporter-specific cost disturbances. Given this condition, σ and α are point identified provided sufficient cross-sectional heterogeneity exists in the relative variances of the two shocks. However, the orthogonality condition fails precisely when markups vary with competitive conditions. If $\gamma \neq 0$, rival prices enter the price-setting equation and the residual ε_{nt}^S absorbs strategic markup responses. Since rival prices respond to the same destination-level demand conditions that drive ε_{nt}^D , the two structural disturbances become correlated, violating the moment condition on which identification rests.

⁸ [Amity et al. \(2019\)](#) employed cost shocks from imported intermediate inputs, while [Gagliardone et al. \(2025\)](#) exploited lagged variation in technical efficiency.

⁹ [Bergstrand et al. \(2023\)](#) derive the same upward-sloping export supply relation from a Melitz model with increasing marginal market-penetration costs, a feature that is nested in our framework through the marginal cost elasticity ϕ , though their framework retains the constant-markup restriction.

¹⁰ Unlike standard LIML (e.g. [Broda and Weinstein, 2006](#); [Imbs and Mejean, 2015](#); [Soderbery, 2015](#); [Grant and Soderbery, 2024](#)), which selects a single reference unit, this difference operator pools all N observations as reference, following [Brasch et al. \(2025\)](#).

3.2 Augmenting the System with Rival Aggregates

To obtain identification in the presence of strategic pricing, we augment the system with rival price and quantity aggregates and recast Equations (17) and (18) in estimable form. Using the expenditure identity $V_{nt} = P_{nt}X_{nt}$ to substitute $\Delta \ln X_{nt}$ out of the demand equation in Equation (17) yields

$$\Delta \ln V_{nt} = \beta \Delta \ln P_{nt} + \sigma \Delta \ln P_{nt}^F + \Delta \ln X_{nt}^F + \Delta e_{nt}^D, \quad (20)$$

where Δe_{nt}^D is a differenced error term. Applying the same substitution to the supply curve in Equation (18) gives

$$\Delta \ln P_{nt} = \Delta \eta_{nt} + \phi (\Delta \ln V_{nt} - \Delta \ln P_{nt}) + \gamma \Delta \ln P_{nt}^F, \quad (21)$$

which rearranges to

$$\Delta \ln P_{nt} = \alpha \Delta \ln V_{nt} + \varsigma \Delta \ln P_{nt}^F + \Delta e_{nt}^S, \quad (22)$$

where Δe_{nt}^S is a differenced supply-side disturbance that captures the idiosyncratic component of the composite intercept $\Delta \eta_{nt}$. The composite parameters map to the structural parameters from Section 2 (Φ and Γ) as

$$\begin{aligned} \alpha &= \frac{\Phi}{1 + \Gamma + \Phi}, \\ \varsigma &= \frac{\Gamma}{1 + \Gamma + \Phi}, \end{aligned} \quad (23)$$

or equivalently, the underlying structural parameters can be recovered via

$$\begin{aligned} \Phi &= \frac{\alpha}{1 - \alpha - \varsigma}, & \phi &= \frac{\alpha}{1 - \alpha}, \\ \Gamma &= \frac{\varsigma}{1 - \alpha - \varsigma}, & \gamma &= \frac{\varsigma}{1 - \alpha}. \end{aligned} \quad (24)$$

3.3 First Stage Estimation: GMM

To estimate the system made up of Equations (20) and (22), we decompose the composite error terms into exporter and time fixed effects plus idiosyncratic disturbances:

$$\begin{aligned} e_{nt}^D &= \lambda_n^D + \lambda_t^D + \varepsilon_{nt}^D, \\ e_{nt}^S &= \lambda_n^S + \lambda_t^S + \varepsilon_{nt}^S. \end{aligned} \quad (25)$$

We identify the parameters by first removing fixed effects in each equation by applying the operator $\ddot{\Delta}$, defined in Section 3.1, to Equations (20) and (22):

$$\begin{aligned} \ddot{\Delta} \ln V_{nt} &= \beta \ddot{\Delta} \ln P_{nt} + F_{nt} \tilde{\beta} + \ddot{\Delta} \varepsilon_{nt}^D, \\ \ddot{\Delta} \ln P_{nt} &= \alpha \ddot{\Delta} \ln V_{nt} + F_{nt} \tilde{\alpha} + \ddot{\Delta} \varepsilon_{nt}^S, \end{aligned} \quad (26)$$

where $F_{nt} = [\ddot{\Delta} \ln P_{nt}^F, \ddot{\Delta} \ln X_{nt}^F]$, $\tilde{\alpha} = [\varsigma, 0]'$, and $\tilde{\beta} = [\sigma, 1]'$.

Stacking the observations over n and t allows us to express the system in matrix form:

$$\begin{aligned} \ddot{\Delta} \ln V &= \beta \ddot{\Delta} \ln P + F \tilde{\beta} + \ddot{\Delta} \varepsilon^D, \\ \ddot{\Delta} \ln P &= \alpha \ddot{\Delta} \ln V + F \tilde{\alpha} + \ddot{\Delta} \varepsilon^S. \end{aligned} \quad (27)$$

In the baseline framework of Section 3.1, F is omitted. If $\gamma \neq 0$, that omission loads both structural disturbances on a common omitted component: rival quantities shift demand, while rival prices enter firms' pricing

decisions. The induced correlation violates the orthogonality condition underlying heteroskedasticity-based identification. We therefore partial out F using the Frisch-Waugh-Lovell theorem. Under the maintained assumption that the dependence between the disturbances operates through these observable rival-market aggregates, the residualized system satisfies the orthogonality condition required for first-stage identification.

We define the projection matrix $\mathcal{P} = F(F'F)^{-1}F'$ and its orthogonal complement $\mathcal{M} = I - \mathcal{P}$, implying $\mathcal{M}F = 0$. Pre-multiplying the system in Equation (27) by $\mathcal{M}$ yields:

$$\begin{aligned}\mathcal{M}\ddot{\Delta} \ln V &= \beta \mathcal{M}\ddot{\Delta} \ln P + \mathcal{M}\ddot{\Delta} \varepsilon^D, \\ \mathcal{M}\ddot{\Delta} \ln P &= \alpha \mathcal{M}\ddot{\Delta} \ln V + \mathcal{M}\ddot{\Delta} \varepsilon^S.\end{aligned}\tag{28}$$

The partialling out step requires $F'F$ to be nonsingular. If F exhibits little cross-sectional variation, $\mathcal{P}$ approaches the identity and $\mathcal{M}$ approaches the zero matrix, causing the rank condition for GMM identification to fail. As established in Section 2.4, this condition holds because exporter-specific destination shares ensure that exporters with heterogeneous destination profiles face distinct competitive conditions, generating genuine cross-sectional variation in F .

Let v_{nt} and p_{nt} denote the nt -th element of the vectors $\mathcal{M}\ddot{\Delta} \ln V$ and $\mathcal{M}\ddot{\Delta} \ln P$, respectively. Similarly, let $\tilde{\varepsilon}_{nt}^D$ and $\tilde{\varepsilon}_{nt}^S$ be the nt -th element of the vectors of transformed residuals $\mathcal{M}\ddot{\Delta} \varepsilon^D$ and $\mathcal{M}\ddot{\Delta} \varepsilon^S$, respectively. Define the scalar variables:

$$Y_{nt} = p_{nt}^2, \quad X_{1nt} = v_{nt}^2, \quad X_{2nt} = p_{nt}v_{nt}, \quad U_{nt} = -\tilde{\varepsilon}_{nt}^D\tilde{\varepsilon}_{nt}^S.\tag{29}$$

Assumption 1 (Identifying Assumptions). *The mean-zero error terms $\tilde{\varepsilon}_{nt}^D$ and $\tilde{\varepsilon}_{nt}^S$ are mutually uncorrelated for any t , such that $\mathbb{E}[\tilde{\varepsilon}_{nt}^D\tilde{\varepsilon}_{nt}^S] = 0$.*

Proposition 1 provides the primary result regarding identification.

Proposition 1 (Moment Equations for Identification). *Consider the econometric framework defined in Equations (26)–(29). Under Assumption 1, we obtain the equation:*

$$Y_{nt} = \theta_1 X_{1nt} + \theta_2 X_{2nt} + U_{nt},\tag{30}$$

where

$$\theta_1 = -\frac{\alpha}{\beta}, \quad \theta_2 = \frac{1}{\beta} + \alpha,$$

and the composite parameter vector $\theta = [\theta_1, \theta_2]'$ is identified from the moment conditions:

$$\mathbb{E}\left[\sum_{t=1}^T U_{nt}\right] = 0 \quad \text{for all } n.\tag{31}$$

Proof. See [Brasch et al. \(2025\)](#). □

Under Assumption 1, GMM estimation of Equation (30) based on the moment conditions in Equation (31) is consistent. Following [Brasch et al. \(2025\)](#), we employ a two-step GMM estimator with an optimal feasible weight matrix in the second step, denoting the resulting estimate as $\hat{\theta}$. The estimator requires $0 \leq \alpha \leq 1$ and $\beta < 0$. Since $\theta_1 = -\alpha/\beta$, this implies $\theta_1 \geq 0$. If $\theta_1 > 0$, the condition $\alpha \leq 1$ is equivalent to $\theta_1 + \theta_2 \leq 1$. Under these conditions, α^{-1} and β are the real solutions to the quadratic equation $\theta_1 s^2 + \theta_2 s - 1 = 0$, yielding the following expressions:

$$\alpha^{-1} = \frac{-\theta_2 + \sqrt{\theta_2^2 + 4\theta_1}}{2\theta_1} > 0, \quad \beta = \frac{-\theta_2 - \sqrt{\theta_2^2 + 4\theta_1}}{2\theta_1} < 0.\tag{32}$$

3.4 Second Stage Estimation and Inference

Since ϕ and γ depend on ς , which was partialled out in the first stage, we recover ς by OLS on the supply curve in Equation (26), requiring the following assumption.

Assumption 2 (Cross-Sectional Independence). *The idiosyncratic disturbances ε_{nt}^S and ε_{nt}^D are uncorrelated across exporters for any t : $\mathbb{E}[\varepsilon_{mt}^S \varepsilon_{nt}^S] = 0$, $\mathbb{E}[\varepsilon_{mt}^D \varepsilon_{nt}^S] = 0$, and $\mathbb{E}[\varepsilon_{mt}^D \varepsilon_{nt}^D] = 0$ for all $m \neq n$.*

Assumption 2 provides the orthogonality condition needed to recover ς from the supply curve in Equation (26). The two assumptions therefore play distinct roles. Assumption 1 is sufficient for the first-stage GMM estimator to recover σ , whereas Assumption 2 is invoked only in the second-stage recovery of ς , and hence of $\gamma = \varsigma/(1 - \alpha)$, as in Equation (24). Thus, when the export price elasticity is the object of interest, identification rests on the orthogonality of demand and supply disturbances in the residualized system in Equation (28). The rival price and quantity aggregates are central to this residualization: absent these controls, strategic pricing would load observable rival-market conditions into the structural disturbances, weakening the basis for Assumption 1 and for the resulting heteroskedasticity-based identification of σ .

Under Assumptions 1 and 2, $\mathbb{E}[\ddot{\Delta} \ln P_{nt}^F \cdot \ddot{\Delta} \varepsilon_{nt}^S] = 0$ holds and the OLS estimator of ς in the rearranged version of the supply curve in Equation (26)

$$\ddot{\Delta} \ln P_{nt} - \hat{\alpha} \ddot{\Delta} \ln V_{nt} = \varsigma \ddot{\Delta} \ln P_{nt}^F + \ddot{\Delta} \varepsilon_{nt}^S \quad (33)$$

is consistent, given by

$$\hat{\varsigma}^{(u)} = \hat{\pi}_1 - \hat{\alpha} \hat{\pi}_2, \quad (34)$$

where $\hat{\pi} = [\hat{\pi}_1, \hat{\pi}_2]$ is the OLS estimator of π in

$$[\ddot{\Delta} \ln P_{nt}, \ddot{\Delta} \ln V_{nt}] = \ddot{\Delta} \ln P_{nt}^F \pi + \text{error}. \quad (35)$$

The structural restrictions $\Gamma \geq 0$ and $\Phi \geq 0$ imply $0 \leq \varsigma \leq 1 - \alpha$. Analogously to the C-GMM estimator, the constrained estimator is:

$$\hat{\varsigma} = \arg \min_{0 \leq \varsigma \leq 1 - \hat{\alpha}} (\varsigma - \hat{\varsigma}^{(u)})^2. \quad (36)$$

The remaining structural parameters are recovered as:

$$\begin{aligned} \hat{\phi} &= \frac{\hat{\alpha}}{1 - \hat{\alpha}}, \\ \hat{\gamma} &= \frac{\hat{\varsigma}}{1 - \hat{\alpha}}. \end{aligned} \quad (37)$$

The two boundary solutions are $\gamma = 0$ and $\phi = \alpha/(1 - \alpha)$ when $\varsigma = 0$, or $\gamma = 1$ and $\phi = \alpha/(1 - \alpha)$ when $\varsigma = 1 - \alpha$.

Standard errors and confidence intervals for all parameters are computed directly from bootstrap draws, without requiring an analytic variance formula. In each iteration $b = 1, \dots, B$, we apply C-GMM to obtain $\hat{\theta}^b$ and then Equations (34) and (36) to obtain $\hat{\varsigma}^b$.¹¹ In the empirical application in Section 4, this is a product-level block bootstrap which preserves within-product serial dependence but does not impose an arbitrary parametric structure on the joint distribution of the product-level moments. The procedure is consistent under the maintained product-level orthogonality condition. However, destination-year shocks and

¹¹ Bootstrap iterations where the unconstrained estimates of θ_1 and θ_2 are inadmissible are discarded.

exchange-rate movements may generate residual dependence across product groups, such that the estimates should be interpreted as conditional on the year effects and rival-market aggregates absorbing the dominant common components of such shocks.

Assumptions 1 and 2 may be violated through two channels. First, an idiosyncratic shock to exporter n may mechanically enter exporter m 's rival aggregates when n has a non-negligible share of their common destination-product market. This channel is attenuated when exporters have small market shares. Second, shocks common to a destination-year may jointly affect the outcomes of several exporters, inducing cross-sectional dependence even when individual market shares are small. We assess the empirical relevance of these channels in the empirical application.

4 Data and Estimation Results

The empirical analysis is based on Norwegian manufacturing exports by product and destination. We construct a panel by product and destination over 1999–2024, combining information on export values, quantities, and prices with competitor prices and quantities from global trade data. We use two sources of data to construct the trade flows: transaction-level trade data from Statistics Norway, and bilateral trade flows data from the CEPII-BACI database.

4.1 Details on Data

4.1.1 Environment

In this section, we focus on a single exporter (Norway) (the index n is fixed and suppressed) and use variation across products (g) and destination markets (d) to identify the model's key parameters. Therefore, the identifying variation comes from differences across products in their destination composition and in the competitor-price and competitor-quantity movements faced in those destinations. Products that sell disproportionately to destinations where competitors' prices increase therefore experience different rival-price shocks than products selling to other destinations. The orthogonality conditions in Section 3 should accordingly be read as restrictions on product-level demand and supply disturbances, after removing product and year effects and after conditioning on the rival price and quantity aggregates.

Following this convention, Norwegian bilateral exports are defined as $(V_{dgt}, X_{dgt}, P_{dgt})$, where the unit value is $P_{dgt} = V_{dgt}/X_{dgt}$. For each destination-product-year cell (d, g, t) , competitor price and quantity indices are constructed from all exporters $j \neq \text{NO}$ observed in the CEPII-BACI data.¹² Aggregating across destinations using Norway's destination-specific sales weights then yields the exporter-level objects $(V_{gt}, X_{gt}, P_{gt}, P_{gt}^F, X_{gt}^F)$ used in the estimation.

4.1.2 Transaction-Level Data From Statistics Norway

Our analysis uses transaction-level data on Norway's foreign trade provided by Statistics Norway. The dataset is based on customs declarations collected through the Norwegian Customs' electronic system for the exchange of declarations (TVINN). Each observation corresponds to a customs declaration line, representing

¹²Excluding Norway from the foreign aggregates corresponds to a country-level small-open-economy assumption: Norway does not affect the foreign price and quantity aggregates, although Norwegian prices may still respond to them. In the small-open-economy definition of [Caliendo and Feenstra \(2024\)](#), a country can affect the relative price of its own variety without affecting relative prices in the rest of the world.

a firm-product-destination combination at the five-digit Standard International Trade Classification (SITC) level.¹³ This granularity allows us to construct firm-level and product-level panels that are essential for estimating export price elasticities.

The coverage of the dataset follows the general trade principle, including all goods that enter or leave Norway’s economic territory, which encompasses Svalbard, Jan Mayen, the continental shelf, and Norwegian vessels and aircraft.¹⁴ Registration occurs when goods physically cross the border. The dataset contains rich information on trade values, quantities, and partner countries. For each transaction, we observe the statistical value, the SITC classification of the good, quantity, and weight. The statistical value represents the free-on-board (FOB) value: the value of the good at the Norwegian border plus any costs of transportation there and insurance, measured in nominal NOK. These variables enable the calculation of unit values, which serve as proxies for prices in the absence of transaction-level price data. Firm identifiers (organization numbers) allow linkage to other administrative registers, facilitating controls for firm industry affiliation.

4.1.3 Bilateral Trade Data From CEPII-BACI

The data from CEPII-BACI (Gaulier and Zignago, 2010) is derived from the UN Comtrade database, and consists of annual bilateral trade flows of products classified according to the Harmonized Commodity Description and Coding System (HS). The UN Comtrade database provides data on imports, exports, re-imports and re-exports (in values and quantities) at the HS six-digit level.¹⁵ Based on this database, the dataset from CEPII-BACI contains information on the value (measured in thousands current U.S. dollars) and the quantity (measured in metric tons) of exports for the period 1995–2024, from which we can derive unit values, for each six-digit HS level of products (approximately 5,000 products) and country (approximately 200 countries).¹⁶

4.2 Construction of Sample and Variables

4.2.1 Sample

From the transaction-level data from Statistics Norway, we discard observations that (i) have an undefined export value, and/or (ii) have missing firm identifiers. We only keep exports of firms in the manufacturing sector. For each good at the five-digit SITC classification level, we calculate the aggregate value and the aggregate quantity for each year. We aggregate the CEPII-BACI trade data from the six-digit HS product level to the five-digit product level based on the SITC classification system. For estimation, we aggregate the data from the five-digit SITC to the three-digit SITC product groups.

Our panel data comprise annual values (V_{dgt}) and quantities (X_{dgt}) of Norwegian exports over the period 1999–2024. We construct unit values P_{dgt} for each observation. For each destination, we also observe competitor prices P_{dgt}^F and quantities X_{dgt}^F .

¹³Until 2007, the SITC classification of goods in the transaction-level data from Statistics Norway was based on Revision 3, while for the period 2007 and onward the classification of goods is based on Revision 4.

¹⁴Note that the transaction-level data from Statistics Norway do not include flows below NOK 1,000.

¹⁵Note that the Comtrade data, and thus the data we use here, do not include flows below USD 1,000.

¹⁶About 8.5 percent of trade quantities in the CEPII-BACI dataset are converted from non-metric units (e.g. meters, watts) into metric tons using product-specific conversion factors. Quantities reported in unknown units or kWh are excluded from the dataset.

Table 1: List of Goods Sections and Export Value

Goods section	SITC Rev. 4 code	Fraction of value
Food and live animals	00–09	17.5
Beverages and tobacco	11–12	0.3
Crude materials, inedible, except fuels	21–29	3.0
Mineral fuels, lubricants and related materials	32–35	2.8
Animal and vegetable oils, fats and waxes	41–43	0.9
Chemicals and related products n.e.s.	51–59	17.1
Manufactured goods classified chiefly by material	61–69	27.1
Machinery and transport equipment	71–79	23.3
Miscellaneous manufactured articles	81–89	7.4
Commodities and transactions	91–97	0.5

Note: Fraction of value refers to the percentage of total export value that each of the 10 goods sections (one-digit SITC level) constitute in total for all goods from the manufacturing industry over the period 1999–2024.

Source: Authors' own calculations using data from Statistics Norway.

4.2.2 Variables

We first use the CEPII-BACI bilateral trade dataset to construct quantities and unit values in levels for each product (g) and year (t). Excluding Norway from the sample (since we are calculating competitor prices and quantities), we construct competitor price and quantity indices for each destination (d) as:

$$\begin{aligned}\Delta \ln P_{dgt}^F &= \sum_{j \neq NO} \psi_{djgt}^F \Delta \ln P_{djgt}, \\ \Delta \ln X_{dgt}^F &= \sum_{j \neq NO} \psi_{djgt}^F \Delta \ln X_{djgt}.\end{aligned}\tag{38}$$

Define the logarithmic mean of two positive numbers a and b as $L(a, b) \equiv (b - a) / (\ln(b) - \ln(a))$ and Sato-Vartia log-mean weights based on competitors' expenditure shares in destination d as:

$$\psi_{djgt}^F = \frac{L(\pi_{djg,t-1}^F, \pi_{djg,t}^F)}{\sum_{k \neq NO} L(\pi_{dkg,t-1}^F, \pi_{dkg,t}^F)}, \text{ where } \pi_{djgt}^F = \frac{P_{djgt} X_{djgt}}{\sum_{k \neq NO} P_{dkgt} X_{dkgt}}.\tag{39}$$

For Norwegian export prices (P_{gt}) and quantities (X_{gt}), we use the indices in Equations (13) and (14) with Sato-Vartia log-mean weights:

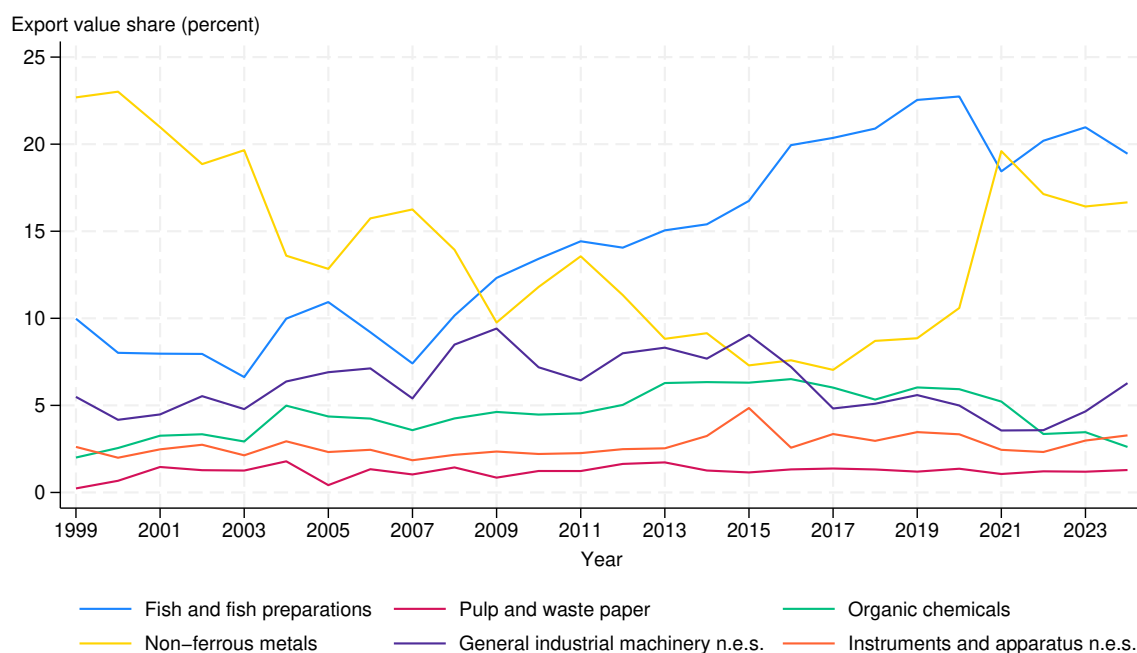
$$\psi_{dgt} = \frac{L(\pi_{dg,t-1}, \pi_{dgt})}{\sum_{m \in \Sigma_{\text{dest}}} L(\pi_{mg,t-1}, \pi_{mgt})}, \text{ where } \pi_{dgt} = \frac{P_{dgt} X_{dgt}}{\sum_{\ell \in \Sigma_{\text{dest}}} P_{\ell gt} X_{\ell gt}}.\tag{40}$$

Competitor price (P_{gt}^F) and volume (X_{gt}^F) indices for each good are constructed using Equations (15) and (16), with weights defined in Equation (40).

4.3 Descriptive Statistics

Table 1 summarizes the product markets in our sample. For expositional clarity, the three-digit SITC goods are aggregated to one-digit SITC sections. Over the sample period, manufactured goods account for 27.1

Figure 1: Export Value Shares, by Year and Largest Goods Divisions



Note: The figure shows the export shares by value from each of the largest goods divisions (two-digit SITC) in each of the six largest goods sections (one-digit SITC) as measured by value shares over the period 1999–2024.

Source: Authors' own calculations using data from Statistics Norway.

percent of total manufacturing export value, while machinery and transport equipment constitute 23.3 percent.

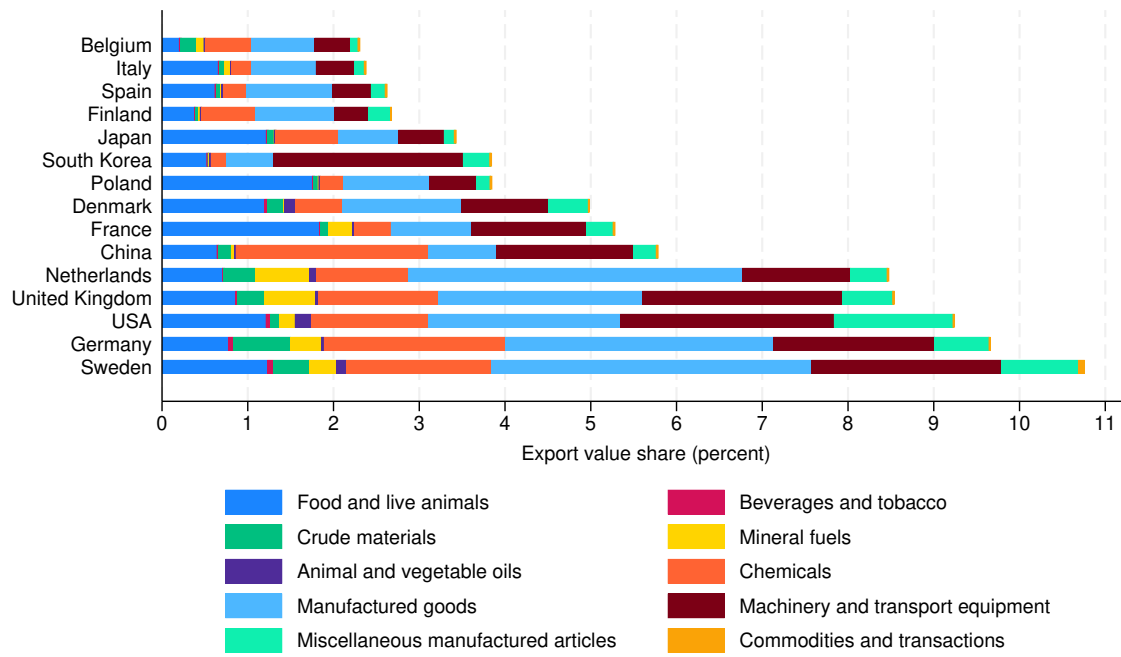
Figure 1 shows the evolution of export-value shares from the largest goods divisions (two-digit SITC) from each of the six largest goods sections (one-digit SITC) as measured by value (as reported in Table 1) over the period 1999–2024. The export of non-ferrous metals dropped from an average of 21 percent of all export value in the manufacturing industry in the period 1999–2003 to 7 percent in 2017, but has since picked up. This product division is consistently the largest in terms of export value through the period we study, with the exception of 2016–2020 when the export of fish and fish preparations was the dominant export good from the manufacturing industry.

Figure 2 shows the distribution of the value of exports from the manufacturing sector to the 15 largest trading partners by goods sections (one-digit SITC), averaged over the sample period. The figure shows that 10.8 percent of all value of exports from the manufacturing sector goes to Sweden, while Germany and the U.S. make up 9.7 and 9.2 percent of exports from the manufacturing industry, respectively.

4.4 Estimation

Table 2 presents the estimation results for the structural and composite parameters for three-digit SITC product groups in Norwegian manufacturing. Because strategic pricing operates through prices as perceived by buyers in the destination market, we express both Norwegian export prices and competitor prices in destination-specific local currency. We convert USD-denominated value amounts from CEPII-BACI to local currency for each destination using IMF IFS exchange rates and re-derive unit values as the ratio

Figure 2: **Export Value Shares, by Destination Country and Goods Section**



Note: The figure shows the 15 largest trading partners in the manufacturing industry (measured by total value share of exports over the period 1999–2024) and the associated export value shares by goods section (one-digit SITC).

Source: Authors' own calculations using data from Statistics Norway.

of local-currency value to quantity.¹⁷ We apply the same conversion to the Norwegian transaction-level data, converting NOK value amounts to destination-specific local currency.

The export price elasticity is estimated to 4.42 with a bootstrap standard error of 0.72. The strategic complementarity elasticity is estimated to 0.51 (standard error of 0.15), which is significantly different from zero at the 1 percent level. The positive estimate indicates that Norwegian exporters adjust prices in response to competitor pricing rather than setting constant markups. The marginal cost elasticity is positive but imprecisely estimated, consistent with modest diminishing returns to scale in production.

The composite parameters are also tightly estimated. The estimate of α reported in Table 2 is 0.57 (standard error of 0.10), indicating a precisely estimated reduced-form feedback from export values to prices in the supply equation. The estimate of ς is 0.22 (standard error of 0.06), implying a statistically significant composite pass-through from competitor prices to own prices, consistent with the positive estimate of the strategic complementarity elasticity γ .

Table A.1 in the Appendix reports additional estimation results for the first-stage GMM parameters θ_1 and θ_2 , the moment-equation coefficients from Equation (30) from which the structural parameters in Table 2 are recovered, together with the underlying structural parameters Γ and Φ . The estimate of the parameter θ_1 is statistically significant at the 1 percent level, while the estimate of θ_2 is statistically significant at the 10 percent level under the normal-approximation test. The estimates of Γ and Φ exhibit large standard errors, reflecting the amplification of sampling uncertainty through the nonlinear mapping in Equation (24) when the denominator $1 - \alpha - \varsigma$ is small.

¹⁷For country-year observations where no official exchange rate is available from IMF International Financial Statistics or World Bank WDI, we retain USD-denominated unit prices.

Table 2: **Estimation Results**

	Parameter Estimate	Bootstrap Standard Error	95 Percent Bootstrap Confidence Interval
<u>Key Parameters</u>			
Elasticity of Demand (σ)	4.42***	0.72	[2.99, 5.84]
Strategic Complementarity Elasticity (γ)	0.51***	0.15	[0.28, 0.91]
Marginal Cost Elasticity (ϕ)	1.34	1.36	[0.67, 5.21]
<u>Composite Parameters</u>			
α	0.57***	0.10	[0.40, 0.84]
ζ	0.22***	0.06	[0.09, 0.33]
Observations ($N \times T$)	6,007		

Note: The parameters α and ζ are composite parameters defined in Equation (23). Sample consists of product groups (three-digit SITC) from Norwegian manufacturing exports, 1999–2024. Competitor prices and Norwegian export prices are expressed in destination-specific local currency. Bootstrap standard errors are based on 498 replications. 2 bootstrap replications were inadmissible ($\hat{\theta}_1^b + \hat{\theta}_2^b > 1$) and discarded. Brackets report bootstrap 95 percent confidence intervals. Significance stars are based on two-sided normal-approximation p -values computed as the point estimate divided by the bootstrap standard error. *** $p < 0.01$.

Source: Authors' own calculations using data from Statistics Norway.

Tables B.1–B.2 in Appendix B assess the sensitivity of the estimates to the construction of the competitor price and quantity indices. The baseline indices use contemporaneous expenditure shares, which may be correlated with the structural shocks in Assumption 1. We therefore report estimates using (i) lagged expenditure shares, (ii) shares fixed at their initial-period values, and (iii) Montgomery-Vartia (Vartia I) weights in place of the baseline Sato-Vartia weights. The point estimate of σ ranges from 4.42 to 4.73 across these variants. The point estimate of γ ranges from 0.29 to 0.51 and remains statistically distinguishable from zero in all but the initial-period variant, where the reweighting that identifies γ is largely eliminated.

We next examine two potential violations of Assumption 2: feedback from Norwegian outcomes to the foreign aggregates, and common shocks that generate price comovement independently of strategic pricing. Table B.3 re-estimates the model after excluding destination-product-year observations in which Norway's market share exceeds thresholds ranging from 20 to 1 percent. Even at the most stringent threshold, only 83 of the 6,007 baseline observations are dropped. As the threshold tightens, $\hat{\sigma}$ declines from 4.42 to 3.41, whereas $\hat{\gamma}$ increases from 0.51 to 0.69 and remains statistically distinguishable from zero throughout. We conclude that the positive estimate of γ is not driven by the small number of destination-product-year cells in which Norway has a relatively large market share. Table B.4 addresses the second channel by controlling for Norway's destination-weighted exposure to destination-year inflation. Local demand or cost conditions specific to a destination-year may raise both Norwegian and competitor prices in that market, generating comovement even in the absence of strategic price adjustment. The resulting estimates, $\hat{\sigma} = 4.41$ and $\hat{\gamma} = 0.51$ (standard error 0.15), are virtually identical to their baseline counterparts.

Beyond these identification concerns, Table B.5 examines whether the results are sensitive to the inclusion of primary-commodity product groups, such as oil and gas, seafood, metals, and ores, whose prices are often set in integrated world markets. For such goods, Norwegian and competitor prices may move together simply because both track a common world price, independently of any strategic price response. Excluding these

groups yields $\hat{\sigma} = 3.59$ and $\hat{\gamma} = 0.46$ (standard error 0.14), both positive and statistically distinguishable from zero. The positive estimate of strategic complementarity is therefore not driven by the inclusion of primary-commodity product groups.

To place our estimates in context, we compare them with prior estimates for Norway and other small open economies. First, our estimate $\hat{\sigma} = 4.42$ considerably exceeds the values embedded in Norway’s main macroeconomic policy models at Norges Bank, Statistics Norway and the Ministry of Finance, where σ ranges from 0.5–1.5; see [Kravik and Mimir \(2019\)](#), [Boug et al. \(2023\)](#) and [Aursland et al. \(2020\)](#). Our estimate also exceeds the aggregate estimate of $\hat{\sigma} = 1.2$ for machinery exports reported in [Boug and Fagereng \(2010\)](#). However, it falls within the upper range of the confidence interval 1.8–5.0 for σ reported by [Imbs and Mejean \(2010\)](#), using Norwegian microdata and a heteroskedasticity-based adaptation of the estimator of [Feenstra \(1994\)](#) that omits strategic pricing. The gap between macro- and micro-based estimates is consistent with [Imbs and Mejean \(2015\)](#), who show that aggregate estimates may understate elasticities when they mask sectoral heterogeneity. [Bergstrand et al. \(2023\)](#) similarly find relatively large trade elasticities when extending the Feenstra framework to allow for firm heterogeneity and increasing marginal costs.

Second, our estimate $\hat{\gamma} = 0.51$ is in line with estimates for small open economies: [Boug et al. \(2006\)](#) report $\hat{\gamma} = 0.64$ for Norwegian machinery using aggregate time-series data, while [Amiti et al. \(2019\)](#) report estimates of γ in the range 0.4–0.6 for Belgium, using microdata and external instruments. [Gagliardone et al. \(2025\)](#) estimate a closely related structural markup elasticity of 0.57 for Belgian manufacturing, which implies a similar degree of strategic complementarity in price setting. Our marginal cost elasticity is positive (but not statistically significant under the normal-approximation test) and is qualitatively consistent with models in which increasing marginal costs arise from market-penetration frictions ([Arkolakis, 2010](#)).

5 Conclusion

Standard heteroskedasticity-based estimators of the export price elasticity rest on a constant-markup assumption that is violated when firms set prices strategically. We develop the first estimator in this class to accommodate strategic pricing. It augments the moment conditions with rival price and quantity aggregates and uses the Frisch-Waugh-Lovell theorem within a constrained GMM framework to restore identification. The estimator jointly identifies the export price elasticity and the strategic complementarity elasticity without requiring quasi-experimental variation or data beyond standard bilateral trade flows and exchange rates.

Using Norwegian manufacturing export data from 1999–2024, we estimate an export price elasticity of 4.42, with a bootstrap standard error of 0.72, and a strategic complementarity elasticity of 0.51, with a bootstrap standard error of 0.15. The precisely estimated complementarity elasticity rejects the constant-markup restriction imposed by standard estimators, indicating that Norwegian exporters adjust prices in response to competitor pricing. The marginal cost elasticity is positive (but not statistically significant), consistent with modest diminishing returns to scale in production.

These findings suggest that allowing for endogenous markup adjustment matters for both identification and interpretation in empirical trade models. More broadly, the framework provides a tractable way to combine heteroskedasticity-based identification with models of variable markups, and may be useful in other applications where strategic pricing is present and valid external instruments are difficult to obtain.

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Appendix

A Additional Estimation Results

Table A.1 shows additional estimation results for the parameters not shown in Table 2.

Table A.1: **Additional Estimation Results**

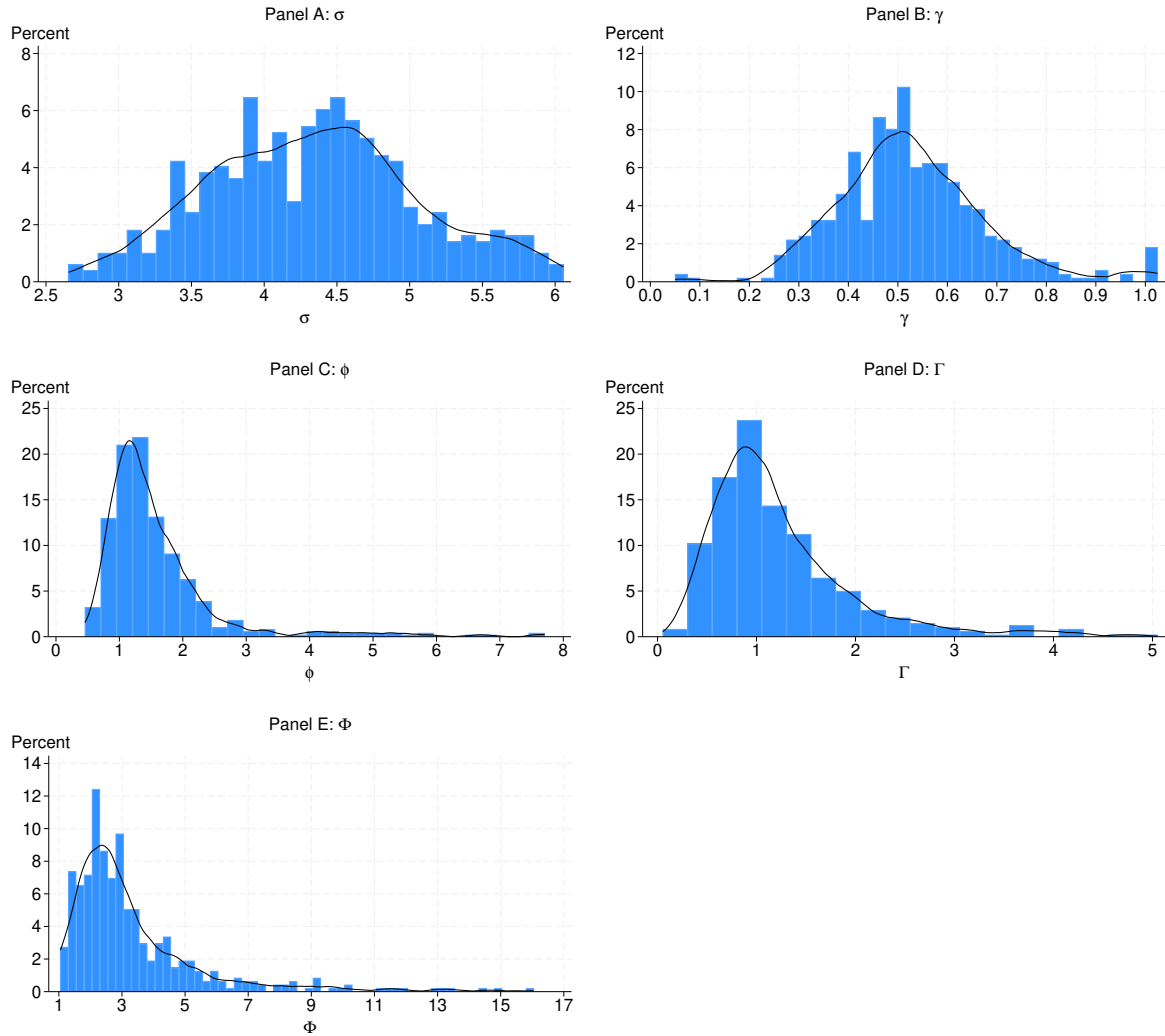
	Parameter Estimate	Bootstrap Standard Error	95 Percent Bootstrap Confidence Interval
<u>GMM Moment-Equation Parameters</u>			
θ_1	0.167***	0.036	[0.121, 0.264]
θ_2	0.280*	0.155	[-0.067, 0.588]
<u>Underlying Structural Parameters</u>			
Γ	1.05	2.14	[0.38, 4.16]
Φ	2.74	10.10	[1.15, 15.73]
Observations ($N \times T$)		6,007	

Note: θ_1 and θ_2 are the GMM moment-equation coefficients from Equation (30), while Γ and Φ are underlying structural parameters recovered via Equation (24). Specification includes exchange-rate indices. Sample consists of product groups (three-digit SITC) from Norwegian manufacturing exports, 1999–2024. Bootstrap standard errors are based on 498 replications. 2 bootstrap replications were inadmissible ($\hat{\theta}_1^b + \hat{\theta}_2^b > 1$) and discarded. Brackets report bootstrap 95 percent confidence intervals. Significance stars are based on two-sided normal-approximation p -values computed as the point estimate divided by the bootstrap standard error. * $p < 0.10$, *** $p < 0.01$.

Source: Authors' own calculations using data from Statistics Norway.

The panels in Figure A.1 show the bootstrap distributions of the estimates for σ , γ , φ , Γ and Φ . The panels are based on 500 bootstrap replications and illustrate the sampling distributions implied by our estimator. For the key structural parameters σ and γ , the bootstrap distributions are tightly centered around the point estimates and approximately symmetric, supporting the use of bootstrap standard errors and percentile confidence intervals. In contrast, the bootstrap distribution of φ is more dispersed, and the distributions of Γ and Φ are even wider, reflecting the amplification of sampling variability through the nonlinear mapping in Equation (24) when the denominator $1 - \alpha - \varsigma$ is small.

Figure A.1: **Distribution of Bootstrap Estimates for Parameters (σ , γ , ϕ , Γ , Φ)**



Note: The figure shows the distribution of the 498 bootstrap estimates for the parameters σ (Panel A), γ (Panel B), ϕ (Panel C), Γ (Panel D) and Φ (Panel E), with a kernel density plot overlaid. 2 bootstrap replications were inadmissible ($\hat{\theta}_1^b + \hat{\theta}_2^b > 1$) and discarded.

Source: Authors' own calculations using data from Statistics Norway.

B Robustness Checks

Table B.1 reports results from two variants of the baseline that replace the contemporaneous Sato-Vartia weights in Equation (38) with predetermined expenditure shares: Panel A uses the expenditure shares (π) from period $t - 1$ rather than t , and Panel B uses shares from the initial period ($\underline{t}$), yielding a geometric Young index where the fixed initial-period shares serve a role analogous to a shift-share (Bartik) design.¹⁸ The motivation is econometric: contemporaneous weights depend on current expenditure shares that may be correlated with the structural shocks in Assumption 1. When fixing the weights at $\underline{t}$, the Sato-Vartia indices in Equations (39) and (40) reduce to the expenditure share in $\underline{t}$:

$$\begin{aligned}\psi_{djgt}^F &= \pi_{djgt}^F = \frac{P_{djgt} X_{djgt}}{\sum_{k \neq NO} P_{dkgt} X_{dkgt}} \\ \psi_{dgt} &= \pi_{dgt} = \frac{P_{dgt} X_{dgt}}{\sum_{\ell \in \Sigma_{dest}} P_{\ell gt} X_{\ell gt}}\end{aligned}\quad (B.1)$$

The two key parameters σ and γ draw their identification from distinct sources of variation, which explains why fixing weights at $\underline{t}$ leaves σ precisely estimated while rendering the estimate of γ insignificant. The export price elasticity σ is identified in the first stage from cross-sectional heteroskedasticity in the structural shocks, a property of the shock distributions that is robust to the choice of weighting scheme. The strategic complementarity elasticity γ , by contrast, is identified in the second stage from cross-sectional variation in competitor prices across goods groups, generated by the dynamic reweighting of destination-specific conditions as trade patterns evolve. Fixing weights at $\underline{t}$ as in Panel B of Table B.1 eliminates this reweighting: rival aggregates reduce to linear combinations of destination-level shocks with time-invariant coefficients, so goods with heterogeneous destination portfolios face near-identical competitive conditions over time, reducing the identifying variation in $F'F$ and attenuating $\hat{\zeta}$. Since $\hat{\gamma} = \hat{\zeta}/(1 - \hat{\alpha})$, this imprecision is amplified by the small denominator ($1 - \hat{\alpha} \approx 0.15$), widening the confidence interval for γ to $[0.00, 1.00]$.

Table B.2 shows estimation of the full system in Equations (20) and (22), but where the Sato-Vartia (Vartia II) weights are replaced with the Montgomery-Vartia (Vartia I) weights. These are defined as follows for the competitor price and quantity indices:

$$\psi_{djgt}^{F,MV} = \frac{\frac{V_{djgt}^F - V_{djg,t-1}^F}{\ln V_{djgt}^F - \ln V_{djg,t-1}^F}}{\frac{\sum_{k \neq NO} V_{dkgt}^F - \sum_{k \neq NO} V_{dkg,t-1}^F}{\ln(\sum_{k \neq NO} V_{dkgt}^F) - \ln(\sum_{k \neq NO} V_{dkg,t-1}^F)}}, \text{ where } V_{djgt}^F = P_{djgt} X_{djgt}. \quad (B.2)$$

For Norwegian export prices and quantity indices, the Montgomery-Vartia (Vartia I) weights are defined as:

$$\psi_{dgt}^{MV} = \frac{\frac{V_{dgt} - V_{dg,t-1}}{\ln V_{dgt} - \ln V_{dg,t-1}}}{\frac{\sum_{m \in \Sigma_d} V_{mgt} - \sum_{m \in \Sigma_d} V_{mg,t-1}}{\ln(\sum_{m \in \Sigma_d} V_{mgt}) - \ln(\sum_{m \in \Sigma_d} V_{mg,t-1})}}, \text{ where } V_{dgt} = P_{dgt} X_{dgt}. \quad (B.3)$$

Table B.3 shows the estimation when destination-good-year cells where Norwegian market share of value ($\bar{s}$) exceeds 20, 15, 10, 5 and 1 percent are dropped from the sample. As is evident from the results, the estimates of the parameters are remarkably stable across all the panels, indicating that the extent to which there is bilateral feedback from markets in which Norway has a non-negligible destination-product market share is negligible in our setting.

¹⁸Panel A requires shares also in $t - 2$, reducing the sample period to 2000–2024.

Table B.4 shows the results from estimating a specification where we include a control for common destination-year shocks that may jointly affect Norwegian and competitor prices even when Norway is a small exporter in each destination-product market. To do so, we control for the destination-year inflation rate, measured as the change in the consumer price index (CPI). For each product-year, we define the inflation-exposure variable as the weighted sum across all destinations:

$$Z_{gt}^{CPI} = \sum_d \pi_{dg,t-1} \pi_{dt}^{CPI}, \quad (\text{B.4})$$

where $\pi_{dt}^{CPI} = \Delta \ln(CPI_{dt})$ and the previous-period weights $(\pi_{dg,t-1})$ are defined in Equation (40). The inflation-exposure variable is orthogonalized against the price and quantity indices, such that:

$$Z_{gt}^{CPI,res} = Z_{gt}^{CPI} - \hat{a} - \hat{b}_{res}^{CPI} \Delta \ln P_{gt}^F - \hat{c}_{res}^{CPI} \Delta \ln X_{gt}^F, \quad (\text{B.5})$$

where $Z_{gt}^{CPI,res}$ is defined as the residual from a linear regression of Z_{gt}^{CPI} on $\Delta \ln P_{gt}^F$ and $\Delta \ln X_{gt}^F$. This residual measure of CPI is then demeaned with regards to the cross-sectional mean ($\tilde{Z}_{gt}^{CPI,res} \equiv Z_{gt}^{CPI,res} - \bar{Z}_{.t}^{CPI,res}$) and included as a variable in Equation (27).

Table B.1: Robustness Check: Weights Based on Different Time Periods

	Parameter Estimate	Bootstrap Standard Error	95 Percent Bootstrap Confidence Interval
A: Expenditure Shares from Previous Period			
<u>Key Parameters</u>			
Elasticity of Demand (σ)	4.58***	0.52	[3.43, 5.53]
Strategic Complementarity Elasticity (γ)	0.47**	0.19	[0.16, 0.95]
Marginal Cost Elasticity (φ)	1.39*	0.77	[0.52, 3.37]
<u>Composite Parameters</u>			
α	0.58***	0.11	[0.34, 0.77]
ς	0.19***	0.07	[0.06, 0.33]
Observations ($N \times T$)		5,769	
B: Expenditure Shares from Initial Period			
<u>Key Parameters</u>			
Elasticity of Demand (σ)	4.73***	0.62	[2.90, 5.38]
Strategic Complementarity Elasticity (γ)	0.30	0.31	[0.00, 1.00]
Marginal Cost Elasticity (φ)	5.89	311.64	[0.71, 103.87]
<u>Composite Parameters</u>			
α	0.85***	0.16	[0.41, 0.99]
ς	0.04	0.07	[0.00, 0.23]
Observations ($N \times T$)		6,002	

Note: Panel A shows the specification where the expenditure shares (π) are from the previous period ($t - 1$), while Panel B shows the specification where π is from the initial period ($\underline{t}$), and thereby also ψ (see Equation (B.1) for more information). The parameters α and ς are composite parameters defined in Equation (23). Sample consists of product groups (three-digit SITC) from Norwegian manufacturing exports, 1999–2024. Competitor prices and Norwegian export prices are expressed in destination-specific local currency. Bootstrap standard errors are based on 500 replications (413 replications in Panel B). 87 bootstrap replications in Panel B were inadmissible ($\hat{\theta}_1^b + \hat{\theta}_2^b > 1$) and discarded. Brackets report bootstrap 95 percent confidence intervals. Significance stars are based on two-sided normal-approximation p -values computed as the point estimate divided by the bootstrap standard error. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Source: Authors' own calculations using data from Statistics Norway.

Table B.2: **Robustness Check: Weights Based on Montgomery-Vartia (Vartia I)**

	Parameter Estimate	Bootstrap Standard Error	95 Percent Bootstrap Confidence Interval
<u>Key Parameters</u>			
Elasticity of Demand (σ)	4.56***	0.80	[3.16, 6.23]
Strategic Complementarity Elasticity (γ)	0.29***	0.07	[0.15, 0.43]
Marginal Cost Elasticity (ϕ)	0.95**	0.47	[0.47, 2.38]
<u>Composite Parameters</u>			
α	0.48***	0.08	[0.31, 0.70]
ζ	0.15***	0.04	[0.06, 0.23]
Observations ($N \times T$)	6,007		

Note: Results from estimation based on the Montgomery-Vartia weights defined in Equations (B.2) and (B.3). The parameters α and ζ are composite parameters defined in Equation (23). Sample consists of product groups (three-digit SITC) from Norwegian manufacturing exports, 1999–2024. Competitor prices and Norwegian export prices are expressed in destination-specific local currency. Bootstrap standard errors are based on 500 replications. Brackets report bootstrap 95 percent confidence intervals. Significance stars are based on two-sided normal-approximation p -values computed as the point estimate divided by the bootstrap standard error. ** $p < 0.05$, *** $p < 0.01$.

Source: Authors' own calculations using data from Statistics Norway.

Table B.3: Robustness Check: Dropping Destination with Large Market Shares

	Parameter Estimate	Bootstrap Standard Error	95 Percent Bootstrap Confidence Interval
A: Share Less than $\bar{s} = 20$ Percent			
Elasticity of Demand (σ)	3.96***	0.57	[3.05, 5.15]
Strategic Complementarity Elasticity (γ)	0.52***	0.17	[0.31, 1.00]
Marginal Cost Elasticity (φ)	1.28	2.59	[0.65, 5.96]
Observations ($N \times T$)		5,993	
B: Share Less than $\bar{s} = 15$ Percent			
Elasticity of Demand (σ)	3.86***	0.56	[2.98, 5.04]
Strategic Complementarity Elasticity (γ)	0.48***	0.17	[0.27, 1.00]
Marginal Cost Elasticity (φ)	1.24	2.47	[0.65, 6.10]
Observations ($N \times T$)		5,989	
C: Share Less than $\bar{s} = 10$ Percent			
Elasticity of Demand (σ)	3.82***	0.55	[3.01, 5.13]
Strategic Complementarity Elasticity (γ)	0.48***	0.18	[0.25, 1.00]
Marginal Cost Elasticity (φ)	1.27	2.37	[0.68, 6.26]
Observations ($N \times T$)		5,983	
D: Share Less than $\bar{s} = 5$ Percent			
Elasticity of Demand (σ)	3.73***	0.52	[2.92, 4.86]
Strategic Complementarity Elasticity (γ)	0.51***	0.18	[0.27, 1.00]
Marginal Cost Elasticity (φ)	1.39	5.59	[0.73, 5.17]
Observations ($N \times T$)		5,977	
E: Share Less than $\bar{s} = 1$ Percent			
Elasticity of Demand (σ)	3.41***	0.40	[2.67, 4.22]
Strategic Complementarity Elasticity (γ)	0.69***	0.23	[0.30, 1.00]
Marginal Cost Elasticity (φ)	2.68	3.01	[1.06, 10.26]
Observations ($N \times T$)		5,924	

Note: Panels A–E show the specification where destination-good-year cells with more than $\bar{s}$ percent market share of Norwegian exports are dropped. Sample consists of product groups (three-digit SITC) from Norwegian manufacturing exports, 1999–2024. Competitor prices and Norwegian export prices are expressed in destination-specific local currency. Bootstrap standard errors are based on 500 replications (497 replications in Panel D and 498 replications in Panel E). 3 bootstrap replications in Panel D and 2 bootstrap replications in Panel E were inadmissible ($\hat{\theta}_1^b + \hat{\theta}_2^b > 1$) and discarded. Brackets report bootstrap 95 percent confidence intervals. Significance stars are based on two-sided normal-approximation p -values computed as the point estimate divided by the bootstrap standard error. *** $p < 0.01$.

Source: Authors' own calculations using data from Statistics Norway.

Table B.4: **Robustness Check: Controlling for Destination-Year Inflation**

	Parameter Estimate	Bootstrap Standard Error	95 Percent Bootstrap Confidence Interval
<u>Key Parameters</u>			
Elasticity of Demand (σ)	4.41***	0.71	[3.00, 5.78]
Strategic Complementarity Elasticity (γ)	0.51***	0.15	[0.27, 0.91]
Marginal Cost Elasticity (ϕ)	1.32	1.44	[0.66, 5.02]
<u>Composite Parameters</u>			
α	0.56***	0.10	[0.39, 0.83]
ς	0.22***	0.06	[0.09, 0.33]
Observations ($N \times T$)	6,007		

Note: Results from estimation when controlling for destination-specific inflation, as defined in Equation (B.5). The parameters α and ς are composite parameters defined in Equation (23). Sample consists of product groups (three-digit SITC) from Norwegian manufacturing exports, 1999–2024. Competitor prices and Norwegian export prices are expressed in destination-specific local currency. Bootstrap standard errors are based on 498 replications. 2 bootstrap replications were inadmissible ($\hat{\theta}_1^b + \hat{\theta}_2^b > 1$) and discarded. Brackets report bootstrap 95 percent confidence intervals. Significance stars are based on two-sided normal-approximation p -values computed as the point estimate divided by the bootstrap standard error. *** $p < 0.01$.

Source: Authors' own calculations using data from Statistics Norway.

Table B.5: **Robustness Check: Removing Raw Materials**

	Parameter Estimate	Bootstrap Standard Error	95 Percent Bootstrap Confidence Interval
<u>Key Parameters</u>			
Elasticity of Demand (σ)	3.59***	0.57	[2.74, 4.97]
Strategic Complementarity Elasticity (γ)	0.46***	0.14	[0.28, 0.99]
Marginal Cost Elasticity (ϕ)	1.00	1.06	[0.48, 3.94]
<u>Composite Parameters</u>			
α	0.50***	0.11	[0.33, 0.80]
ς	0.23***	0.05	[0.11, 0.32]
Observations ($N \times T$)	5,498		

Note: Results from estimation based where raw material goods are dropped from the estimation sample. The parameters α and ς are composite parameters defined in Equation (23). Sample consists of product groups (three-digit SITC) from Norwegian manufacturing exports, 1999–2024, but excluding (i) oil and gas, (ii) seafood, (iii) metals, (iv) wood and pulp, and (v) ores and minerals. The three-digit SITC for these product groups are: 333, 034, 035, 036, 037, 682, 683, 684, 685, 686, 687, 688, 689, 247, 248, 251, 273, 278, 281, 284, 287 and 288. Competitor prices and Norwegian export prices are expressed in destination-specific local currency. Bootstrap standard errors are based on 500 replications. Brackets report bootstrap 95 percent confidence intervals. Significance stars are based on two-sided normal-approximation p -values computed as the point estimate divided by the bootstrap standard error. *** $p < 0.01$.

Source: Authors' own calculations using data from Statistics Norway.