Venting Out: Exports during a Domestic Slump

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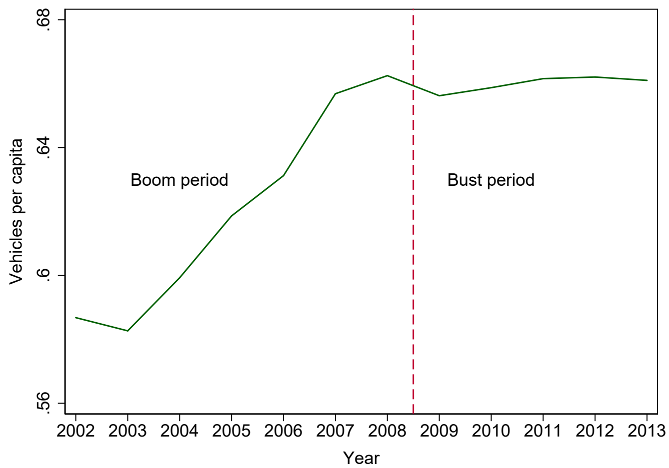
Online Appendix (Not for Publication)

C Appendix Figures

C.1 Evolution of the Stock of Vehicles in Spain

Figure C.1 plots the evolution of the stock of vehicles per capita in Spain over the period 2002-13. The figure illustrates the abrupt stop in 2009 of the expansion in that stock during the boom period. Source: DGT (n.d.).

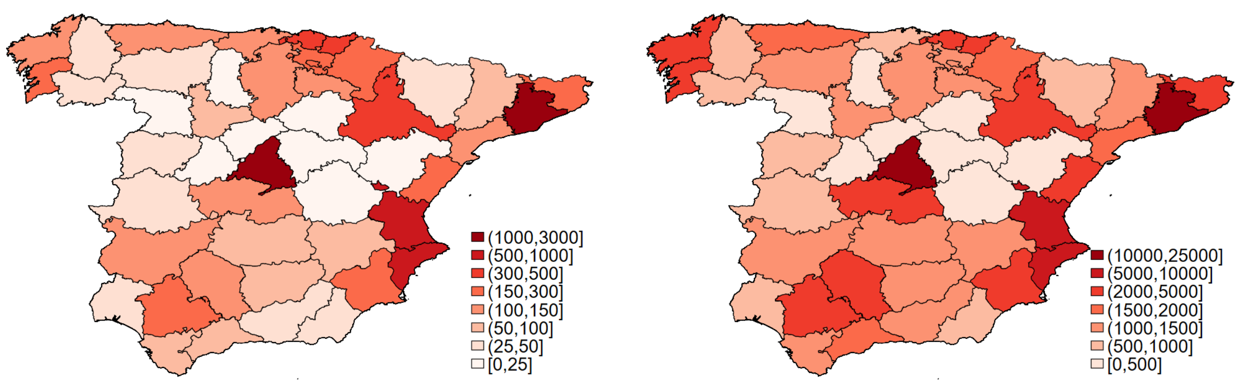
Figure C.1: Stock of Vehicles per Capita in Spain



C.2 Spatial Distribution of Economic Activity in Spain

Figure C.2 plots the 2002-08 yearly average number of firms and number of exporting firms for each of the 47 Spanish peninsular provinces. Sources: (Banco de España, n.d.*a*,n; Informa, n.d.; IGN, n.d.).

Figure C.2: Distribution of Economic Activity in Spain: Variation Across Provinces



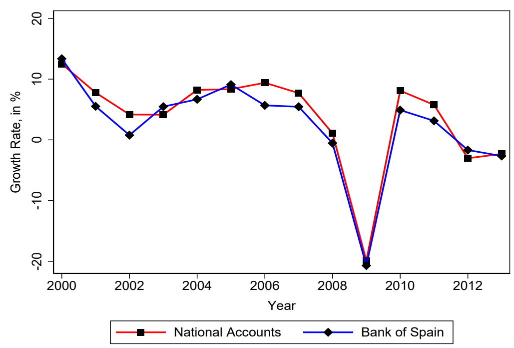
(a) Number of Firms (b) Number of Exporting Firms

C.3 Data Coverage of Macroeconomic Dynamics

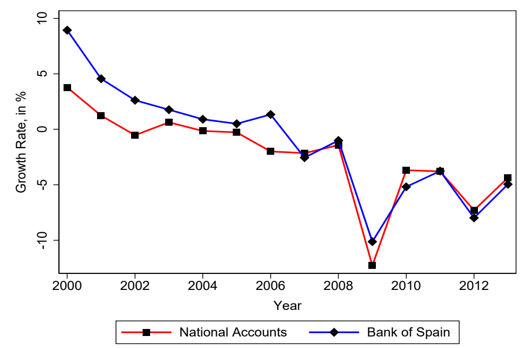
Figure C.3 compares the annual growth rates of key economic variables in our dataset with the annual growth rates reported in official publicly available aggregate data from National Accounts

(INE, n.d.*a*) and Customs (Ministerio de Hacienda, n.d.). The figures show that our dataset tracks well the aggregate evolution over time of output, employment, total payments to labor, and exports.

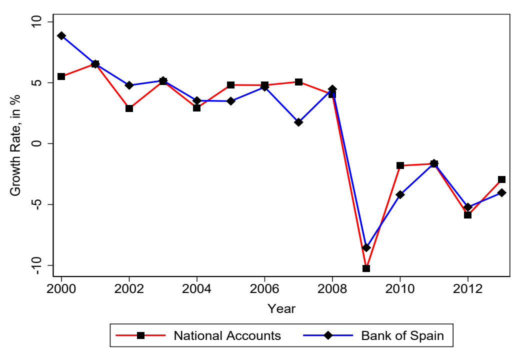
Figure C.3: Output, Employment, Wage Bill and Export Dynamics



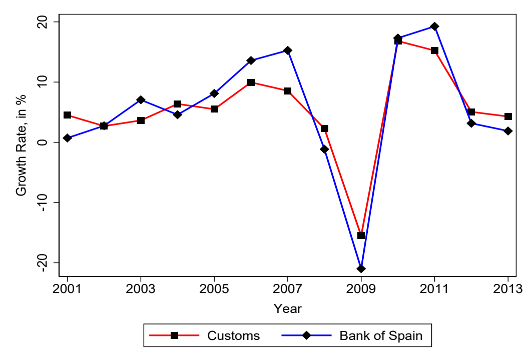
Panel (a): Output



Panel (b): Employment



Panel (c): Wage Bill



Panel (d): Exports

C.4 Variation in Domestic Sales and Vehicles per capita at the Municipal Level

Figure C.4 illustrates variation across zip codes in both the boom-to-bust changes in average manufacturing firm-level domestic sales and the boom-to-bust changes in the number of vehicles per capita. We do so for the case of the two most populated provinces in Spain: Madrid and Barcelona. To facilitate a comparison of the within-province across-zip codes variation illustrated in Figure C.4 with the across-province variation illustrated in Figure 3, the average zip code changes illustrated in Figure C.4 have been standardized using the Spain-wide mean and cross-province standard deviation used to standardize the corresponding variables in Figure 3.

Panels (a) and (b) reveal a large heterogeneity in the change in both firms' average domestic sales and vehicles per capita across zip codes located in the region of Madrid: while the center area of the region that contains a large number of tightly packed zip codes (this area corresponds to the city of Madrid) experienced relatively small reductions in firm average domestic sales, surrounding zip codes experienced changes that were more than two standard deviations above the national average. Similarly, while the zip codes belonging to the city of Madrid experienced a large reduction in the number of vehicles per capita (more than two standard deviations smaller than the Spain-wide average), other zip codes to the east, north and west of the city saw increases in vehicles per capita significantly above the national average. Panels (c) and (d) provide analogous information for the province of Barcelona. Although the heterogeneity across zip codes is smaller than that observed within the Madrid region, panel (c) still shows that certain zip codes experienced growth rates smaller than the national average while others experienced changes in firm average domestic sales more than a standard deviation above that average.

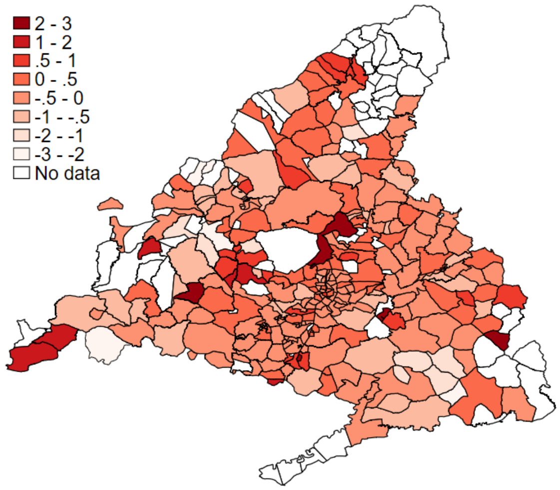
C.5 First-Stage and Reduced-Form Relationships

The two panels in Figure C.5 provide a graphical representation of the relationship between our baseline instruments (i.e., the local instrument, in panel A, and the gravity-based instrument, in panel B) and the boom-to-bust change in both the log of domestic sales (left figures) and exports (right figures). In each panel, left figures thus represent the first-stage relationship between the endogenous covariate and the instrument, while right figures represent the reduced-form relationship between the outcome variable of interest and the instrument.

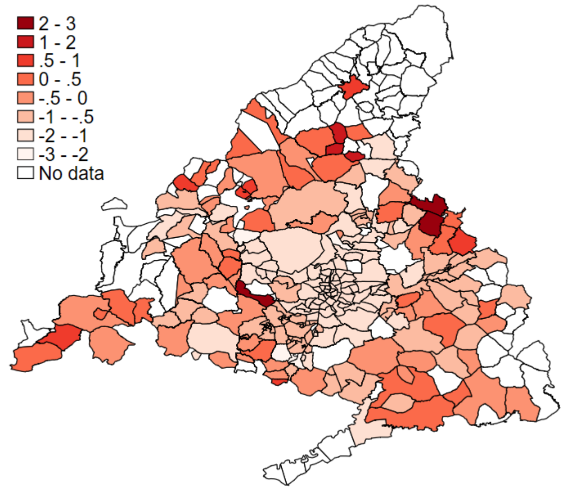
C.6 Variation in the Instrument at the Municipal Level

Figure C.6 shows the distribution of firm-level deviations from provincial means for the local instrument (the change in log vehicles per capita in municipality) and for the gravity-based instrument (the change in log distance-population-weighted vehicles per capita). The variables are normalized such that a value of 1 for a given observation means that it is one standard deviation away from the mean in the province where the firm is located. The histograms show that there is substantial variation in the instruments *within* provinces, although there is naturally a large share of observations within two standard deviations of the provincial means. There are also a few outliers on both sides, as one would expect if the data generating process is a standard normal distribution.

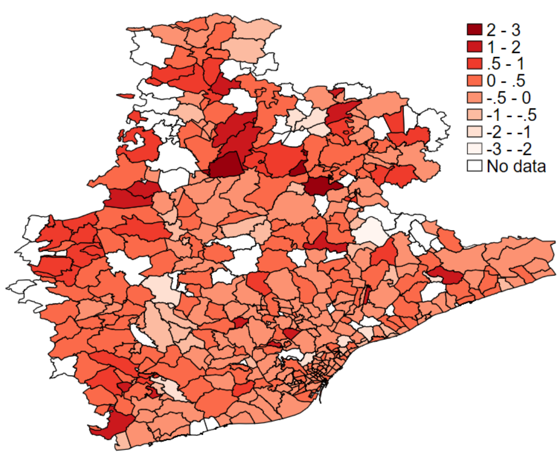
Figure C.4: The Great Recession in Madrid and Barcelona: Variation Across Zip Codes



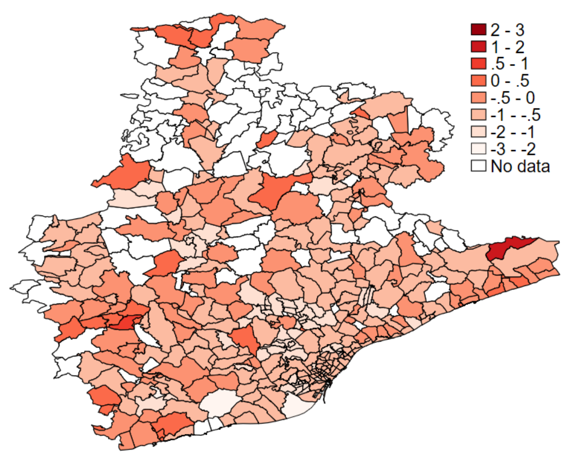
(a) Relative Change in Domestic Sales (Madrid province)



(b) Relative Change in Vehicles per Capita (Madrid province)



(c) Relative Change in Domestic Sales (Barcelona province)



(d) Relative Change in Vehicles per Capita (Barcelona province)

Notes: Panel (a) illustrates the standardized percentage change in average firm-level domestic sales between the period 2002-2008 and the period 2009-2013. Therefore, if this variable takes any given value p for a given zip code, it means that the average firm located in this zip code experienced a relative change in average yearly domestic sales between 2002-2008 and 2009-2013 that was p standard deviations above the change experienced by a firm located in the (Spain-wide) mean zip code. Panel (b) illustrates the standardized percentage change in cars per capita between the period 2002-2008 and the period 2009-2013. Therefore, if this variable takes any given value p for a given zip code, it means that this zip code experienced a relative change in vehicles per capita between 2002-2008 and 2009-2013 that was p standard deviations above the change experienced by the (Spain-wide) mean zip code. Zip codes that do not host any of the firms in our dataset appear in white, with the label “No data”.

D Additional Macroeconomic Evidence

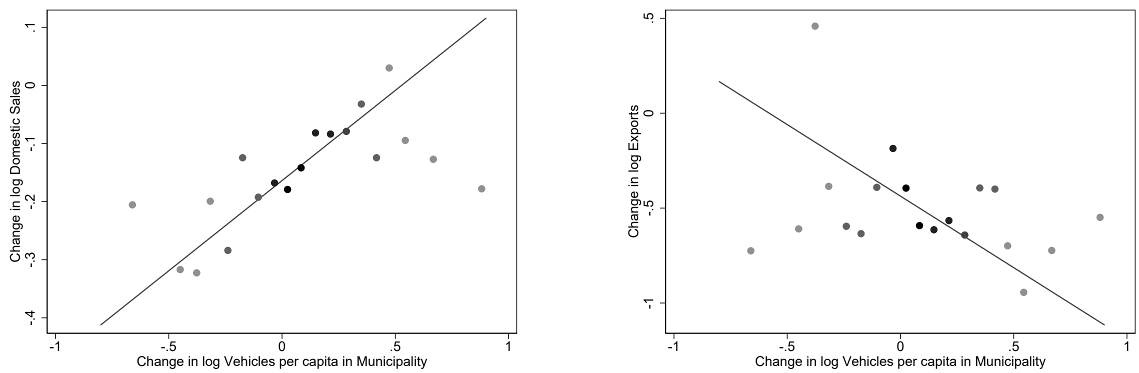
D.1 Basic Motivating Facts for Spain and Other Countries

In this Appendix, we present figures analogous to Figure 1 for a wider set of EMU-12 countries, and explore how the findings would change if we exclude the export of vehicles from the export series.

Each of the figures in this Appendix contains two panels. Panel (a) plots, relative to the total

Figure C.5: First-Stage and Reduced-Form Relationships

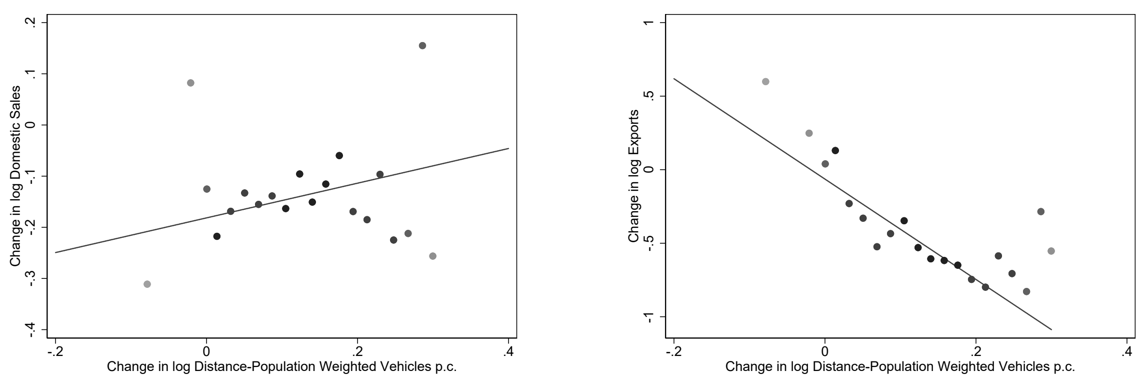
(a) Panel A: Local Instrument



First-Stage Relationship

Reduced-Form Relationship

(b) Panel B: Gravity-based Instrument

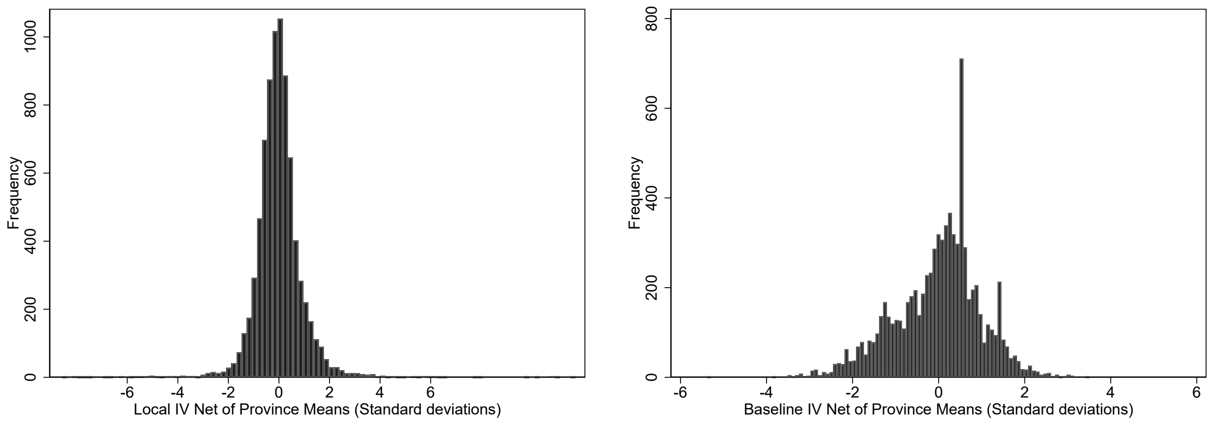


First-Stage Relationship

Reduced-Form Relationship

Notes: Each dot in these figures represents the average change in log domestic sales (left figures) and in log exports (right figures) for a given value of: (a) the local instrument (change in the log of stock of vehicles at the municipal level) in Panel A; or (b) the gravity-based instrument in Panel B, i.e. the change in the municipality-specific distance- and population-weighted average of the stock of vehicles per capita in every other municipality, with weights built using the estimates in column 1 of Table 1. Observations are grouped into 30 equal-sized intervals of the horizontal axis, with the exception of cases where a bin contains five or less observations (which are grouped together to reduce the influence of outliers). The darkness of the markers is proportional to the number of observations in each bin. The regression lines depicted are estimated using the same number of observations (N=8,009) as in the regressions of Table 3, without including any controls or fixed effects.

Figure C.6: Within-Province Variation in the Two Instruments



Note: these figures show the distribution of deviations from province means in the local instrument (change in log vehicles per capita in municipality), on the left panel, and in the gravity-based instrument, on the right panel. In both figure, the variable is normalized, so the horizontal axis measures how far, in standard deviations, each firm-level observation is from the province mean.

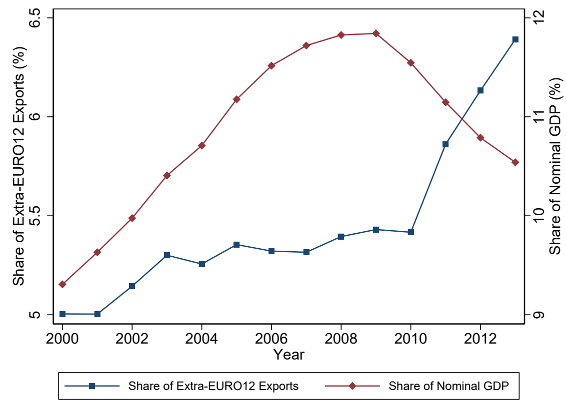
for all EMU-12 countries, a country's share of goods exports to non-EMU-12 countries and its share of nominal GDP. Panel (b) includes an analogous plot but excludes from the export data all export flows in the HS2 category "Vehicles; other than railway or tramway rolling stock, and parts and accessories thereof" (HS2 code 87). The data on exports is from UN Comtrade (United Nations, n.d.), and the nominal GDP data is from the AMECO database. (European Commission, n.d.). We present these figures for eight countries: Spain, Portugal, Greece, Ireland, Italy, Germany, France, and the Netherlands.

Several observations are in order. First, the patterns in Spain, Portugal and Greece are quite similar, with the steep decline in the relative GDP of these countries around the crisis being accompanied by a significant increase in their export share to non-EMU-12 countries. Second, we observe in Germany and France the mirror image of the patterns observed in Southern Europe, with an increase in the relative GDP of those countries and a decline in their export share around the crisis. Third, the cases of the Netherlands and Italy are distinct in that one observes a fairly stable positive correlation between the relative GDP and relative export shares of those countries. Finally, excluding the motor vehicle industry from the relative export series has a minor effect on these figures. This means, in particular, that the Spanish export miracle has little to do with dynamics in that sector.

These figures illustrate that the macroeconomic facts that motivate our study are also relevant to other EMU countries. Whether the vent-for-surplus mechanism was a key factor behind these facts for countries other than Spain is left as an open question for future research.

Figure D.1: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Spain

(a) All Industries



(b) Excluding Motor Vehicle Industry

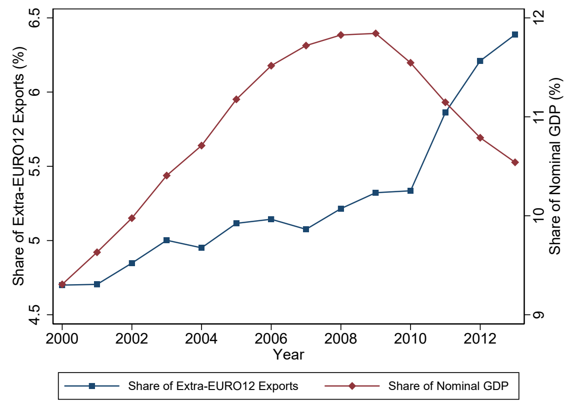
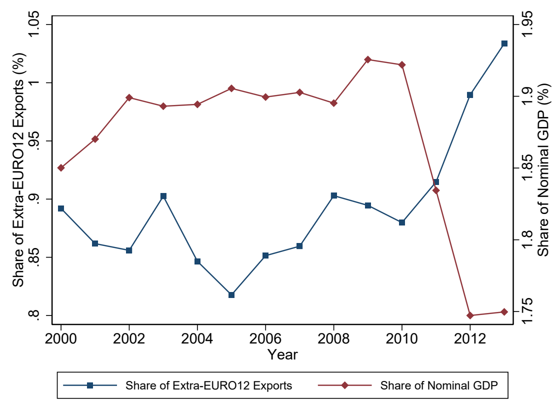


Figure D.2: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Portugal

(a) All Industries



(b) Excluding Motor Vehicle Industry

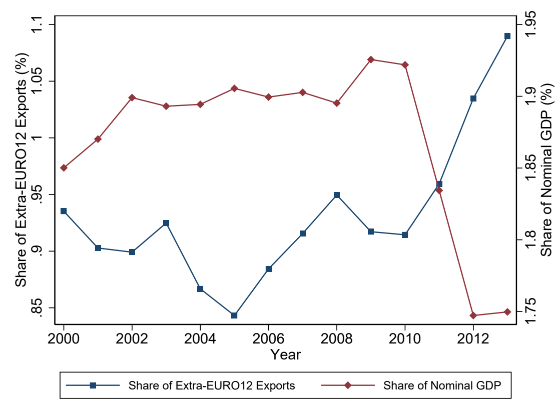
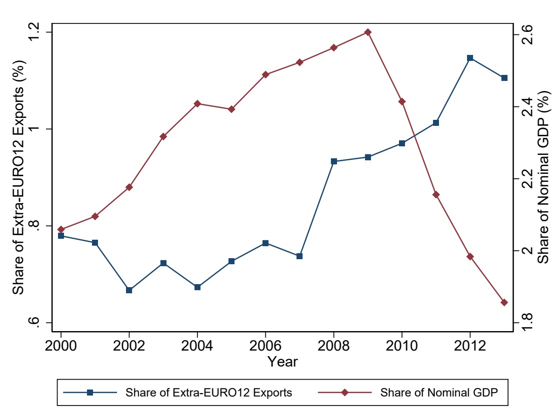


Figure D.3: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Greece

(a) All Industries



(b) Excluding Motor Vehicle Industry

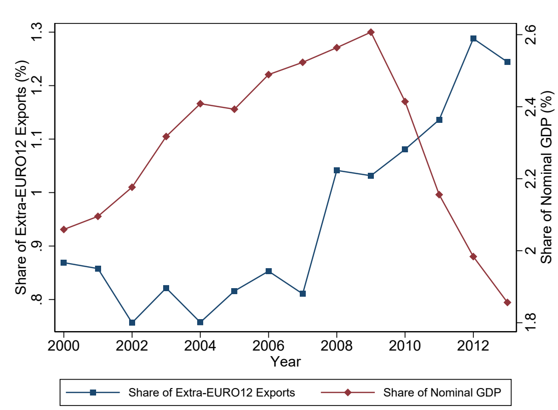


Figure D.4: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Ireland

(a) All Industries (b) Excluding Motor Vehicle Industry

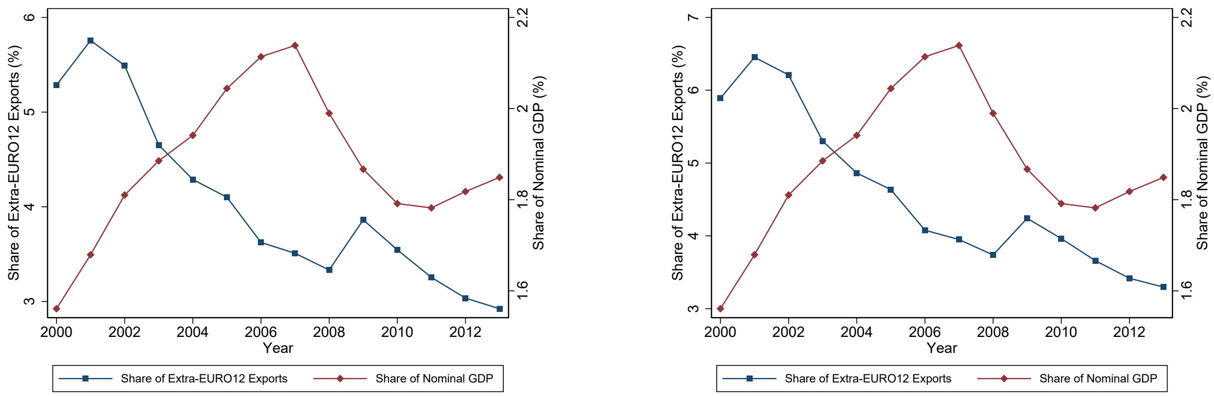


Figure D.5: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Italy

(a) All Industries (b) Excluding Motor Vehicle Industry

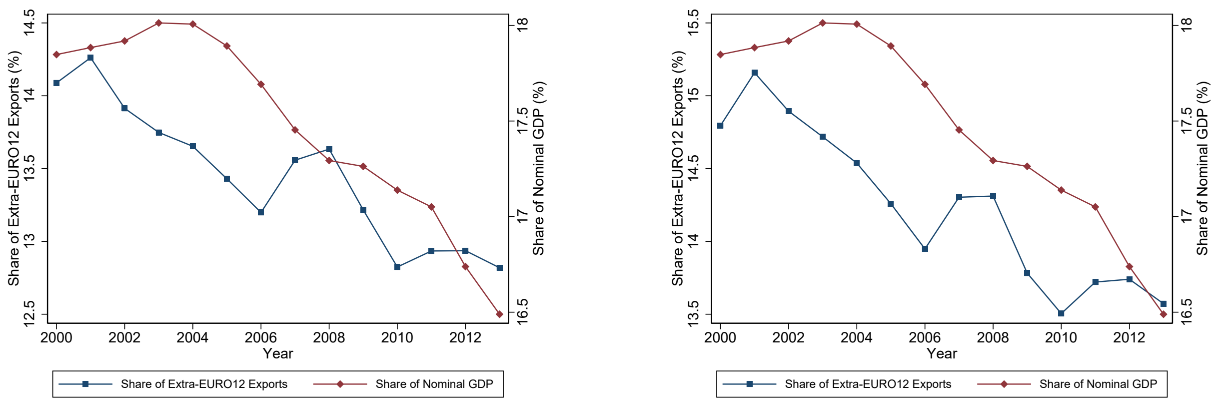


Figure D.6: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Germany

(a) All Industries (b) Excluding Motor Vehicle Industry

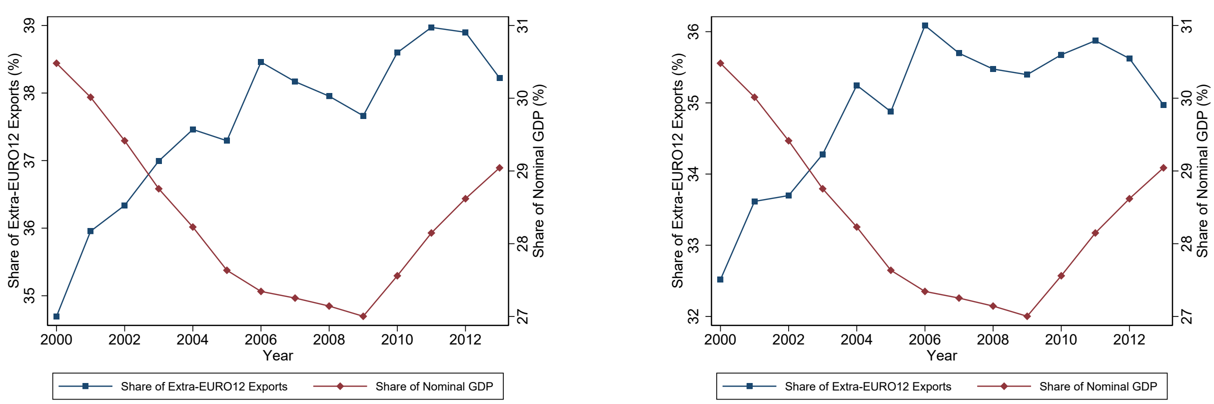


Figure D.7: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: France

(a) All Industries (b) Excluding Motor Vehicle Industry

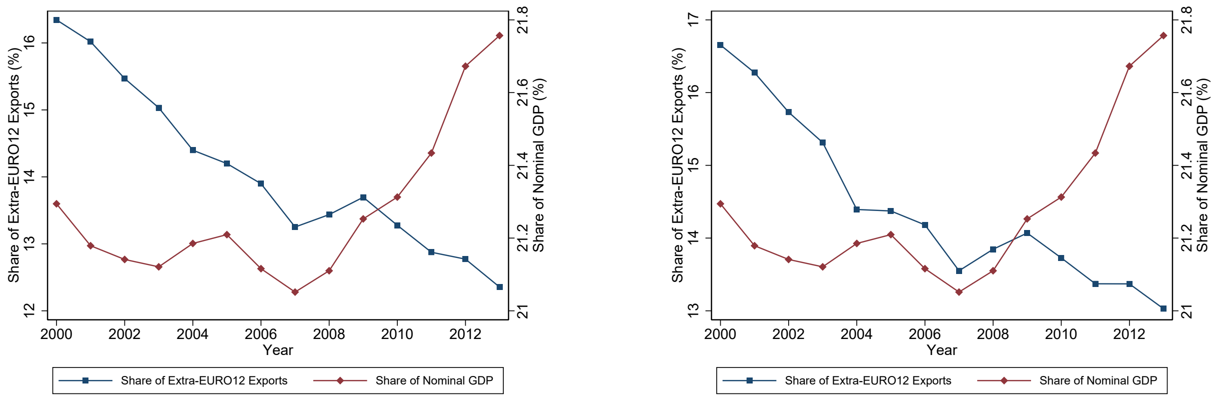
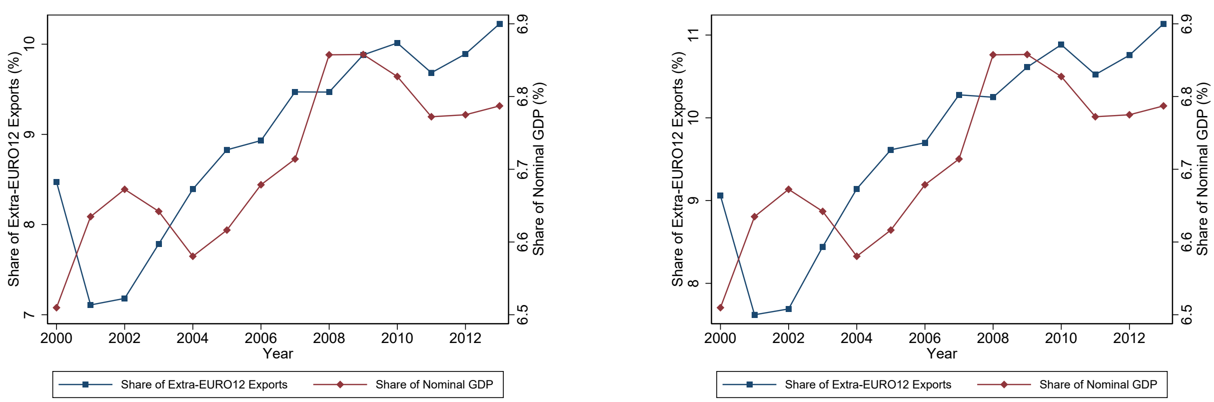


Figure D.8: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Netherlands

(a) All Industries (b) Excluding Motor Vehicle Industry

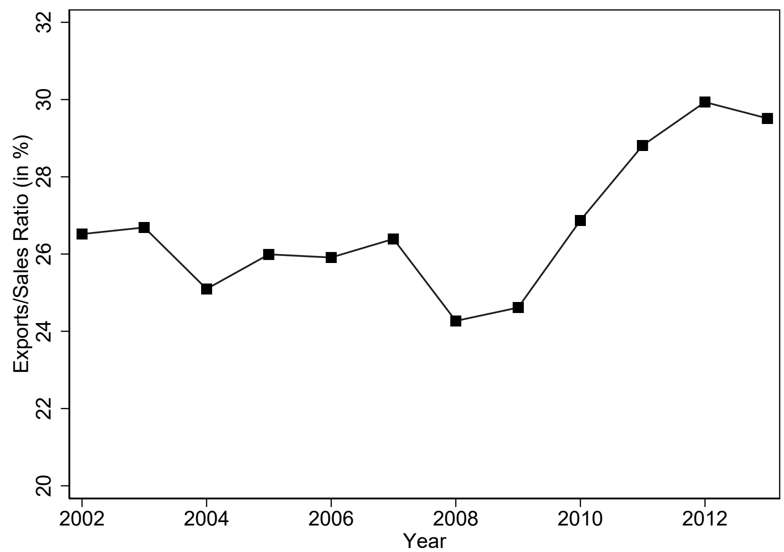


D.2 Behavior of the Exports-to-Sales Ratio in Spain

The dynamics of the exports-to-sales ratio for firms in our baseline sample of continuing exporters is consistent with the macro evidence, as shown in Figure D.9 below. The exports-to-sales ratio is stable around 26% in the boom period (2002-2008) and increases sharply to about 30% during the recession, especially between 2009 and 2012.

There is considerable variation in the exports-to-sales ratio across sectors. We divide the 21 subsectors of manufacturing considered in our analysis (note that we always exclude tobacco, oil refining and motor vehicles) into three groups depending on whether they had negative, low (positive but below 2%) or high (above 2%) growth in their exports-to-sales ratio between the boom and bust periods. The annual patterns are shown in Figure D.10. The ratio is flat or declining in traditional sectors like beverages, textiles, paper, leather and shoes, and also for printing and chemicals. It features moderate growth in sectors like food, clothing, wood, fabricated metals, pharmaceutical products and computers & electronics. Finally, the exports-to-sales ratio increases strongly mostly in sectors such as electrical equipment, machinery & equipment (including repairs), other transportation equipment (different from vehicles), furniture, rubber & plastics, basic metals, and nonmetallic minerals.

Figure D.9: Exports-to-Sales Ratio for Continuing Exporters (Baseline Sample)



D.3 Behavior of Export Prices

In this section, we describe the time series of the unit values of Spanish exports relative to that of the eleven countries (other than Spain) that belonged to the monetary union all throughout our sample period: Austria, Belgium, France, Finland, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands and Portugal. With this aim, we collect data from UN Comtrade (United Nations, n.d.) on the export unit values P_{jdst} for every one of these countries of origin j, every possible destination market d, every HS6 product code s, and every year t between 2002 and 2013. Using this data, we estimate the year fixed effects λ_t and the sector- and destination-specific fixed effects μ_{ds} in the following regression:

$$\ln(P_{SPdst}/P_{EUdst}) = \lambda_t + \mu_{ds} + \varepsilon_{dst}, \tag{D.1}$$

where P_{SPdst} is Spain's export unit value to destination d in good s and year t, P_{EUdst} is a weighted average of the export unit values P_{jdst} across all eleven countries other than Spain belonging to the European Monetary Union for all years between 2002 and 2006, and ε_{dst} is a regression residual.

We estimate the parameters of the regression in equation (D.1) under two alternative definitions of the average price P_{EUdst}. First, we use an average price that uses weights that are fixed over time:

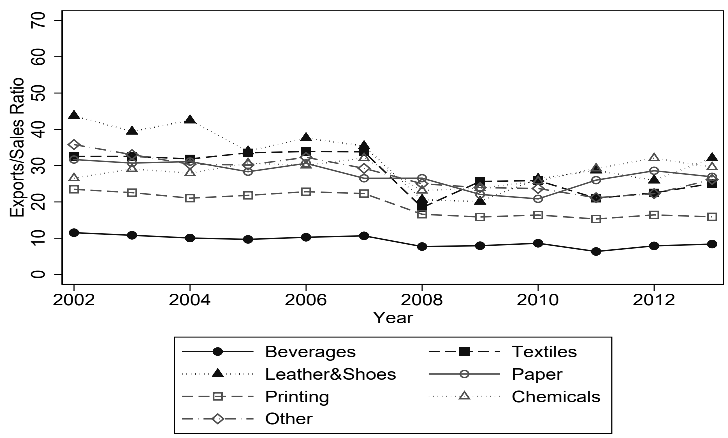
$$P_{EUdst} = \sum_{j \neq SP} \omega_{jds2000} P_{jdst}, \tag{D.2}$$

where $\omega_{jds2000}$ equals the ratio of exports of good s from country j to destination d in the year 2000 relative to the total exports of good s from all eleven countries we have selected as comparison group. Second, we use an average price that uses weights that vary over time:

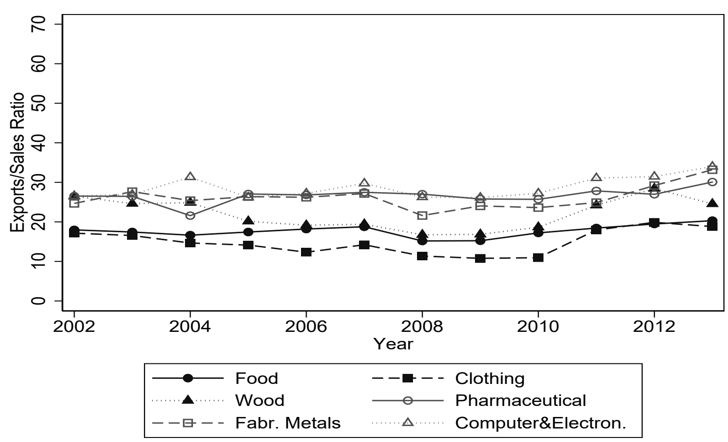
$$P_{EUdst} = \sum_{j \neq SP} \omega_{jdst} P_{jdst}, \tag{D.3}$$

Figure D.10: Exports-to-Sales Ratio by Sector

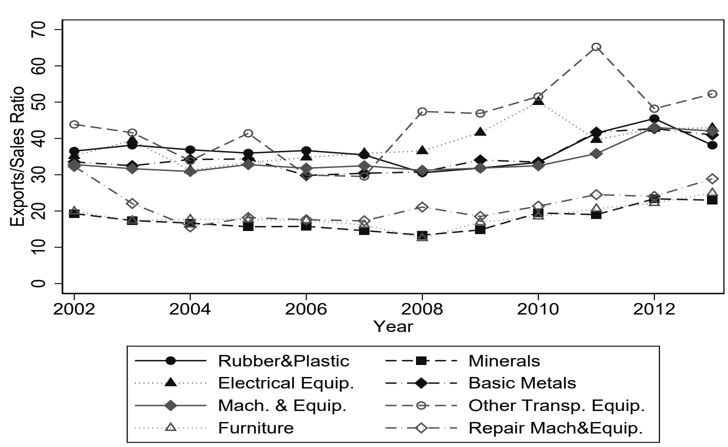
(a) Negative Growth Sectors



(b) Low Growth Sectors



(c) High Growth Sectors

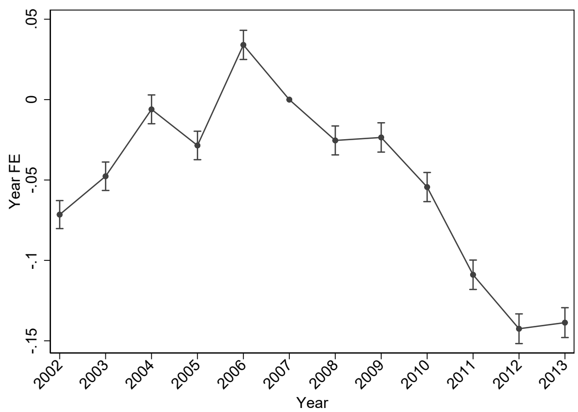


where ω_{jdst} equals the ratio of exports of good s from country j to destination d in year t relative to the total exports of good s from all eleven countries we have selected as comparison group.

Figure D.11 reports the OLS estimates of λ_t in equation (D.1). These estimates are normalized so that $\lambda_{2007} = 0$. Panel (a) reports the corresponding estimates for the case in which P_{EUdst} is computed using the expression in equation (D.2); Panel (b) reports similar estimates when the formula in equation (D.3) is used to compute P_{EUdst}. Both panels illustrate that the unit values of Spanish exports relative to those of the eleven countries (other than Spain) that belonged to the monetary union all throughout our sample period generally increased between 2002 and 2007 and generally decreased between 2008 and 2012, having remained stable between 2012 and 2013.

Figure D.11: Export Unit Values of Spain Relative to Other Euro-area Countries

(a) With Constant Weights



(b) With Time-Varying Weights

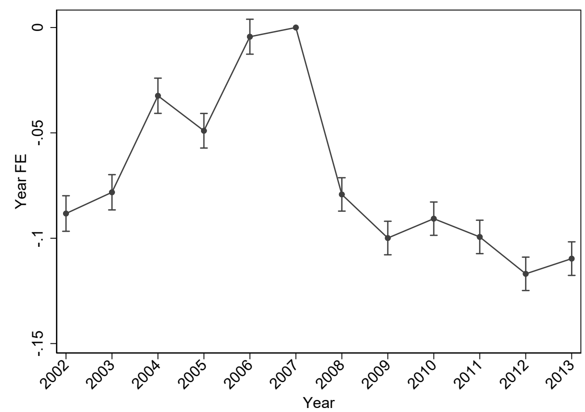
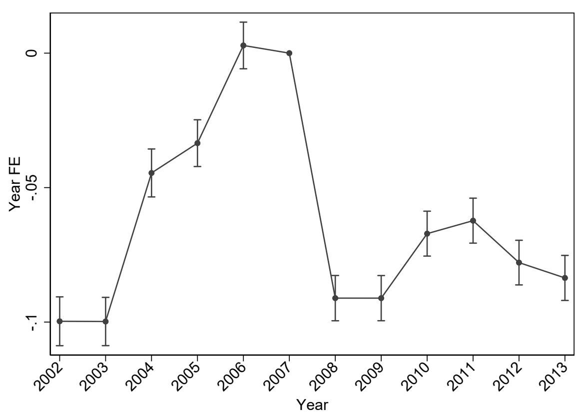
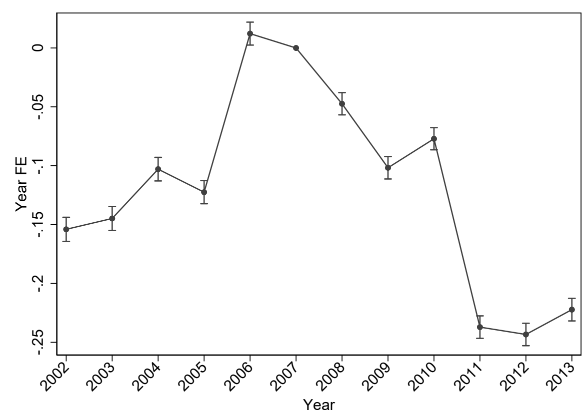


Figure D.12: Export Unit Values of Spain Relative to Germany and France

(a) Relative to Germany



(b) Relative to France



A comparison of Figure D.11 with Figure 1 shows that the evolution of relative Spanish export unit values is positively correlated with the evolution of relative Spanish GDP and negatively correlated with the evolution of relative Spanish exports. These facts are consistent with our model. Because of an increasing marginal cost curve, prices are higher when total production is higher, and vice versa. They are also consistent with the relative reduction in export unit values explaining the extraordinary growth in total exports that took place in Spain (relative to other

Euro countries) during the bust period.

Figure D.12 presents graphs analogous to those in Figure D.11 for the two largest countries in the European Monetary Union. As the two panels of Figure D.12 illustrate, it is in both cases true that Spanish export unit values grew relatively more during the years up to 2007, and decreased relatively more during the post 2007 years.

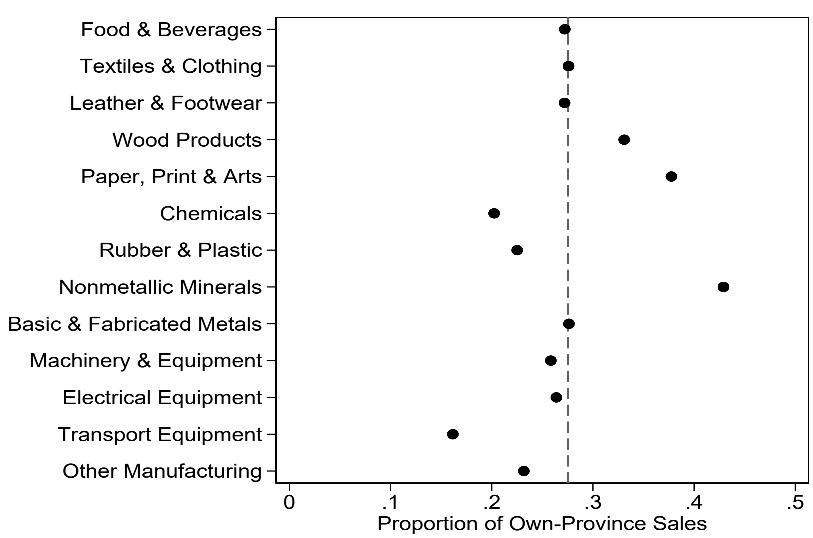
D.4 Home Bias in Firms' Tax Records versus the C-Intereg Dataset

In this Appendix, we briefly compare some aggregate statistics on firms' within-Spain sales from our tax records and from the C-Intereg dataset (CEPREDE, n.d.). We will then provide a sectoral decomposition of the home bias in the latter dataset.

In our 2006 aggregate data on municipality-to-municipality flows for firms in the manufacturing sector (which excludes sales of businesses in the auto industry), the average share of sales in which the origin and destination are the same municipality is 8.75%. This share increases to 25.4% when we consider sales within the seller's own province and to 34.9% for sales in the seller's region (or "Autonomous Community", to use the legal term in Spain).[1]

C-Intereg is a micro-database constructed from a random sample of shipments by road within Spain during the period 2003-2007. Although we do not have access to this micro-database, we have obtained province-to-province shipments from that database.[2] An advantage of the C-Intereg dataset is that it is representative at the sectoral level and, thus, allows to compute valid estimates of sector-specific own-province shares of shipments. As Figure D.13 demonstrates, the own-province sales shares ranges from a low 18% in Transport Equipment (an industry we exclude from our analysis) to a high of 43% for Nonmetallic Minerals. The overall provincial home bias in manufacturing in the C-Intereg dataset is 28%, which is quite close the 25.4% we computed in our dataset based on tax records.

Figure D.13: Province-Level Home Bias in Spanish Manufacturing



[1] In Spain, there are 8,018 municipalities, which are part of 50 provinces, and 17 autonomous communities. There are also two autonomous cities, Ceuta and Melilla.

[2] We are grateful to Carlos Llano for providing them to us; see Llano et al. (2010) and https://www.c-intereg.es/en/ for more details on this database.

E Econometric Biases

E.1 Biases Due to Measurement Error in Exports and Total Sales

We discuss here the implications of measurement error in total sales and exports when a firm's domestic sales are computed by subtracting its exports from its total sales (see also Berman, Berthou and Héricourt, 2015).

Suppose that one does not observe R_{id} directly, but instead measures it as $R_i - R_{ix}$, where R_i and R_{ix} denote the total sales and aggregate exports of firm i, respectively. Assume furthermore that both $\Delta \ln R_i$ and $\Delta \ln R_{ix}$ are measured with error, so that

$$\Delta \ln R_i = \Delta \ln \breve{R}_i + \varpi_i \qquad \text{and} \qquad \Delta \ln R_{ix} = \Delta \ln \breve{R}_{ix} + \varpi_{ix},$$

where $\breve{R}_i$ and $\breve{R}_{ix}$ denote the true values of total sales and exports. Note then that

$$\Delta \ln R_{id} = \Delta \ln R_i - \Delta \ln R_{ix} = \Delta \ln \breve{R}_i - \Delta \ln \breve{R}_{ix} + \varpi_i - \varpi_{ix},$$

Following the same steps as in the main text, we reach an estimating equation that differs from that in equation (6) only in that the regression residual now includes the measurement error in exports; i.e.,

$$\varepsilon_{ix} = (\sigma - 1)\,[u^{\xi}_{ix} + u^{\varphi}_i - u^{\omega}_i] + \varpi_{ix}.$$

Similarly, we reach an expression analogous to that in equation (8) except that the regression residual now depends on the measurement error in total sales and exports; i.e.,

$$\varepsilon_{id} = (\sigma - 1)\,[u^{\xi}_{id} + u^{\varphi}_i - u^{\omega}_i] + \varpi_{iT} - \varpi_{ix}.$$

It then follows that the probability limit of the OLS estimator

$$plim(\hat{\beta}_{OLS}) = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}_{id})}{var(\Delta \ln \mathcal{R}_{id})}$$

can we written as

$$plim(\hat{\beta}_{OLS}) = \frac{cov(u^{\xi}_{ix} + u^{\varphi}_i - u^{\omega}_i + \frac{1}{\sigma-1}\varpi_{ix}, u^{\xi}_{id} + u^{\varphi}_i - u^{\omega}_i + \frac{1}{\sigma-1}\,(\varpi_{iT} - \varpi_{ix}))}{var(u^{\xi}_{id} + u^{\varphi}_i - u^{\omega}_i + \frac{1}{\sigma-1}\,(\varpi_{iT} - \varpi_{ix}))}.$$

This expression is analogous to that in equation (11) but it highlights the potential for additional sources of bias related to the covariance between the measurement error terms ϖ_{ix} and $\varpi_{iT} - \varpi_{ix}$. The sign of this bias depends on the correlation between the measurement errors in total sales and in exports. If these variables are constructed from different sources (e.g., total sales are obtained from census data, while exports are drawn from customs data) it seems plausible that these measurement errors will be uncorrelated with each other; in this case, the impact of measurement errors in total sales and exports on the bias in the OLS estimate $\hat{\beta}_{OLS}$ will be negative. Nevertheless, if the measurement errors in total sales and exports are highly correlated and the variance of the measurement error in total sales is larger than that of the measurement error in exports, it is possible for these measurement errors to contribute to the positive bias in the OLS estimate $\hat{\beta}_{OLS}$.

Consider next an IV estimator of β, where $\Delta \ln R_{id}$ is instrumented with a variable Z_{id}. The probability limit of this IV estimator is

$$plim(\hat{\beta}_{IV}) = \frac{cov(u^{\xi}_{ix} + u^{\varphi}_{i} - u^{\omega}_{i} + \frac{1}{\sigma-1}\varpi_{ix}, \mathcal{Z}_{id})}{cov(u^{\xi}_{id} + u^{\varphi}_{i} - u^{\omega}_{i} + \frac{1}{\sigma-1}\left(\varpi_{iT} - \varpi_{ix}\right), \mathcal{Z}_{id})}.$$

This expression illustrates that $plim(\hat{\beta}_{IV}) = 0$ as long as the instrument Z_{id} verifies three conditions: (a) it is correlated with the change in domestic sales of firm i after partialling out sector fixed effects and the observable determinants of the firm's marginal cost that we include in our regression specification; (b) it is mean independent of the change in firm-specific unobserved productivity, u^{φ}_{i}, factor costs, u^{ω}_{i}, and export demand u^{ξ}_{ix}; and (c) it is mean independent of the measurement error in exports ϖ_{ix}.

E.2 Analysis of the Model with Multiple Domestic and Foreign Markets

The benchmark model described in section VI accounts for one aggregate domestic market and one aggregate foreign market. We generalize here this model to incorporate multiple domestic markets and multiple foreign markets. One can interpret the multiple domestic markets as capturing Spanish municipalities, and the multiple foreign markets as capturing foreign countries.

We exploit the model described here to study how the existence of multiple domestic and foreign markets impacts possible biases affecting the OLS and several IV estimators of the elasticity of a firm's total exports with respect to demand-driven changes in the firm's total sales (i.e., sum of total exports and aggregate domestic sales). Specifically, we present simulation results that are informative about the sign of these biases and their quantitative importance.

E.2.1 Model: Description and Solution Algorithm

Notation. We use i to index firms, j to index markets, and t to index time periods. We use J_{it} to denote the set of markets to which firm i sells at period t. Similarly, we use J_{idt} and J_{ixt} to denote the set of domestic and foreign markets, respectively, to which a firm exports at period t.

Demand function. We assume that the demand in any market j at period t for the variety produced by firm i is

$$Q_{ijt} = \frac{P^{-\sigma}_{ijt}}{P^{1-\sigma}_{jt}} E_{jt} \xi^{\sigma-1}_{ijt}. \tag{E.1}$$

Rearranging terms, we can write the optimal price P_{ijt} as

$$P_{ijt} = Q^{-\frac{1}{\sigma}}_{ijt} P^{\frac{\sigma-1}{\sigma}}_{jt} E^{\frac{1}{\sigma}}_{jt} \xi^{\frac{\sigma-1}{\sigma}}_{ijt}. \tag{E.2}$$

Variable costs. Firm i's total variable cost of producing Q_{ijt} units in each market $j \in J_{it}$ is

$$\frac{1}{\varphi_{it}} \omega_{it} \frac{1}{\lambda+1} (Q_{it})^{\lambda+1} \qquad \text{with} \qquad Q_{it} \equiv \sum_{j \in J_{it}} \tau_{ij} Q_{ijt}.$$

The marginal cost to firm i in period t of selling Q_{ijt} in market j is thus

$$\frac{\omega_{it}\tau_{ij}}{\varphi_{it}} (\sum_{j \in J_{it}} \tau_{ij} Q_{ijt})^{\lambda}. \tag{E.3}$$

Fixed costs. We assume that a firm i has to pay fixed costs F_{ijt} to sell a positive amount in market j at period t.

Market structure. We assume that firms are monopolistically competitive and, thus, their optimal price in any market j at period t is equal to a constant markup $\sigma/(\sigma-1)$ over the marginal cost of selling in market j at t. Therefore, using the marginal cost expression in equation (E.3), the price set by firm i in market j in period t is:

$$P_{ijt} = \frac{\sigma}{\sigma-1}\frac{\omega_{it}\tau_{ij}}{\varphi_{it}}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{\lambda}. \tag{E.4}$$

Market-specific sales. The optimal quantity sold by firm i in each market j belonging to the set J_{it} is determined as the outcome of the following optimization problem

$$\max_{Q_{ijt}\in J_{it}}\Big\{Q_{ijt}^{\frac{\sigma-1}{\sigma}}P_{jt}^{\frac{\sigma-1}{\sigma}}E_{jt}^{\frac{1}{\sigma}}\xi_{ijt}^{\frac{\sigma-1}{\sigma}} - \frac{1}{\varphi_{it}}\omega_{it}\frac{1}{\lambda+1}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{\lambda+1}\Big\}.$$

The first-order condition corresponding to each $Q_{ijt}\in J_{it}$ is

$$\frac{\sigma-1}{\sigma}Q_{ijt}^{-\frac{1}{\sigma}}P_{jt}^{\frac{\sigma-1}{\sigma}}E_{jt}^{\frac{1}{\sigma}}\xi_{ijt}^{\frac{\sigma-1}{\sigma}} - \frac{\omega_{it}\tau_{ij}}{\varphi_{it}}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{\lambda} = 0,$$

and, thus, the optimal quantity sold by firm i in market j at period t is

$$Q_{ijt} = E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{1}{P_{jt}}\Big)^{1-\sigma}\Big(\frac{\sigma}{\sigma-1}\frac{\omega_{it}\tau_{ij}}{\varphi_{it}}\Big)^{-\sigma}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{-\lambda\sigma}. \tag{E.5}$$

Combining equations (E.4) and (E.5), we can write the optimal revenue of firm i in market j at period t as

$$R_{ijt} \equiv Q_{ijt}P_{ijt} = E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\frac{\sigma}{\sigma-1}\frac{\omega_{it}}{\varphi_{it}}\Big)^{-(\sigma-1)}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{-\lambda(\sigma-1)}. \tag{E.6}$$

Furthermore, given that,

$$\begin{aligned}\sum_{j\in J_{it}}\tau_{ij}Q_{ijt} = \sum_{j\in J_{it}}\tau_{ij}\frac{R_{ijt}}{P_{ijt}} &= \sum_{j\in J_{it}}R_{ijt}\frac{\sigma-1}{\sigma}\frac{\varphi_{it}}{\omega_{it}}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{-\lambda}\\ &= \frac{\sigma-1}{\sigma}\frac{\varphi_{it}}{\omega_{it}}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{-\lambda}\sum_{j\in J_{it}}R_{ijt}\\ &= \frac{\sigma-1}{\sigma}\frac{\varphi_{it}}{\omega_{it}}(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{-\lambda}R_{it},\end{aligned}$$

we can rewrite

$$(\sum_{j\in J_{it}}\tau_{ij}Q_{ijt})^{1+\lambda} = \frac{\sigma-1}{\sigma}\frac{\varphi_{it}}{\omega_{it}}R_{it}$$

and

$$(\sum_{j\in J_{it}} \tau_{ij} Q_{ijt})^{-\lambda(\sigma-1)} = \Big(\frac{\sigma-1}{\sigma}\frac{\varphi_{it}}{\omega_{it}} R_{it}\Big)^{-\frac{\lambda(\sigma-1)}{1+\lambda}}.$$

Plugging this expression in equation (E.6), we can further rewrite the optimal revenue of firm i in market j at period t as

$$\begin{aligned} R_{ijt} &= E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\frac{\sigma}{\sigma-1}\frac{\omega_{it}}{\varphi_{it}}\Big)^{-(\sigma-1)}\Big(\frac{\sigma-1}{\sigma}\frac{\varphi_{it}}{\omega_{it}} R_{it}\Big)^{-\frac{\lambda(\sigma-1)}{1+\lambda}} \\ &= E_{jt}\Big(\frac{\sigma}{\sigma-1}\frac{\omega_{it}}{\varphi_{it}}\Big)^{-\frac{\sigma-1}{1+\lambda}}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\Big)^{-(\sigma-1)}\big(R_{it}\big)^{-\frac{\lambda(\sigma-1)}{1+\lambda}}, \end{aligned} \tag{E.7}$$

where R_{it} denotes the total sales of firm i at period t.

Aggregate domestic sales and total exports. Thus, using the expression in equation (E.7), we can write aggregate domestic sales as

$$R_{idt} = \sum_{j\in J_{idt}} R_{ijt} = \Big(\frac{\sigma}{\sigma-1}\frac{\omega_{it}}{\varphi_{it}}\Big)^{-\frac{\sigma-1}{1+\lambda}}\big(R_{it}\big)^{-\frac{\lambda(\sigma-1)}{1+\lambda}}\sum_{j\in J_{idt}} E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\Big)^{-(\sigma-1)}, \tag{E.8}$$

total exports as

$$R_{ixt} = \sum_{j\in J_{ixt}} R_{ijt} = \Big(\frac{\sigma}{\sigma-1}\frac{\omega_{it}}{\varphi_{it}}\Big)^{-\frac{\sigma-1}{1+\lambda}}\big(R_{it}\big)^{-\frac{\lambda(\sigma-1)}{1+\lambda}}\sum_{j\in J_{ixt}} E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\Big)^{-(\sigma-1)}, \tag{E.9}$$

and total sales as

$$R_{it} = R_{idt} + R_{ixt} = \Big[\Big(\frac{\sigma}{\sigma-1}\frac{\omega_{it}}{\varphi_{it}}\Big)^{-\frac{\sigma-1}{1+\lambda}}\sum_{j\in J_{it}} E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\Big)^{-(\sigma-1)}\Big]^{\frac{1+\lambda}{1+\lambda\sigma}}. \tag{E.10}$$

Market participation. The optimal variable profits that a firm i will make in a market j and year t upon entry is given by

$$\pi_{ijt} = R_{ijt} - \frac{1}{\varphi_{it}}\omega_{it}\frac{1}{\lambda+1}\frac{\sigma-1}{\sigma}\frac{\varphi_{it}}{\omega_{it}}R_{it} = \Big(1-\frac{1}{\lambda+1}\frac{\sigma-1}{\sigma}\Big)R_{ijt}. \tag{E.11}$$

Given that a firm has to pay fixed costs F_{ijt} to sell a positive amount in market j at period t, the total profits that a firm i will make in any given market j and year t upon entry is given by

$$\Pi_{ijt} = \pi_{ijt} - F_{ijt}. \tag{E.12}$$

Consequently, the total profits of a firm i selling in a set of markets J_{it} at period t is

$$\Pi_{it}^{J_{it}} = \sum_{j\in J_{it}} \Pi_{ijt}. \tag{E.13}$$

Firms are assumed to select the set of markets they sell to in order to maximize their total profits;

thus, the optimal set of markets a firm sells to at period t is

$$J_{it} = \arg\max_{J} \Pi_{it}^{J}, \tag{E.14}$$

where J denotes a generic set of possible markets to which a firm may sell to.

Solution algorithm. The empirically relevant objects entering our estimating equations are the aggregate domestic sales, R_{idt}, total exports, R_{ixt}, and total sales R_{it}. For every possible firm i and period t, we implement the following procedure in order to compute R_{idt}, R_{ixt} and R_{it}.

Our procedure requires looping over every every possible set J of markets to which firm i may sell to in period t. For each J, we implement the following steps. First, we use equation (E.10) to compute R_{it}. Second, we use equations (E.7), (E.11) and (E.12) to compute R_{ijt}, π_{ijt}, and Π_{ijt}, respectively, for every market j belonging to the set J. Third, we use equation (E.13) to compute Π_{it}^{J}. Once we know Π_{it}^{J} for every possible set of markets J to which firm i may sell to in period t, we use (E.14) to compute the optimal set of markets J_{it} in which firm i sells at period t. Knowledge of J_{it} implies knowledge of J_{idt} and J_{ixt}. Once we know J_{idt}, J_{ixt}, and J_{it}, we use equations (E.8) to (E.10) to compute R_{idt}, R_{ixt} and R_{it}. Our approach becomes computationally infeasible when the number of destinations is large, but we will restrict our simulations to a relatively low number of destinations. As mentioned below, one could use iterative algorithms to scale up our exercise to a larger number of destinations.

E.2.2 Estimating Equations, Estimators and Endogeneity Problems

The parameter of interest in our empirical application is the elasticity of total exports with respect to total sales. To mimic the baseline boom-to-bust identification approach we follow in the main draft, we derive an estimating equation from the model described in Appendix E.2.1 that compares the outcomes between two periods $t = 0$ and $t = 1$.

We use the notation $\Delta \ln(X_{it}) \equiv \ln(X_{i1}) - \ln(X_{i0})$ for every possible variable X and firm i. Then, from equation (E.9), we can write an estimating equation analogous to that in equation (17) in section VI.A as

$$\begin{aligned} \Delta \ln(R_{ixt}) &= -\frac{\sigma - 1}{1+\lambda}\big(\Delta \ln(\omega_{it}) - \Delta \ln(\varphi_{it})\big) - \frac{\lambda(\sigma-1)}{1+\lambda}\Delta \ln(R_{it}) + \Delta \nu_{it}^{x}, \\ \Delta \nu_{it}^{x} &= \Delta \ln\Big(\sum_{j \in J_{ixt}} E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\Big)^{-(\sigma-1)}\Big). \end{aligned} \tag{E.15}$$

The parameter of interest in our empirical application is thus $-\lambda(\sigma-1)/(1+\lambda)$. When estimating this parameter, we treat all variables in this regression equation expect for $\Delta\nu_{it}^{x}$ as observed. We treat the term $\Delta\nu_{it}^{x}$ as unobserved, as it depends on unobserved firm-, market- and year-specific demand shocks ξ_{ijt} and unobserved market and year-specific price indices P_{jt}.[3,4]

In section E.2.3, we illustrate the properties of both the OLS estimator as well as of two instrumental variable (IV) estimators that use each a different instrument for $\Delta \ln(R_{it})$. These two

[3] In our empirical application (see section VI.A), we allow the supply shifters $\Delta \ln(\omega_{it})$ and $\Delta \ln(\varphi_{it})$ to be imperfectly observed. This does not affect qualitatively the properties of the OLS and IV estimators, as the lack of observability of the demand shifter ξ_{ijt} and price index P_{jt} causes possible estimators of the parameter $\lambda(\sigma-1)/(1+\lambda)$ to be biased in a way similar to how they would be if the supply shifters were not perfectly controlled for.

[4] Although the expression in equation (E.15) results from aggregating a destination-specific gravity equation, the issues that arise in the estimation of the parameter of interest are different from those discussed in Redding and Weinstein (2019). The reason is that neither the relevant regressor nor the parameter of interest (i.e., neither $\Delta \ln(R_{it})$ nor $\lambda(\sigma-1)/(1+\lambda)$) vary by export market. Thus, when aggregating across destinations, the relevant regressor appears outside the summation of terms over destinations.

instruments are measures of the market potential of a firm i at period t in the domestic market. Specifically, indexing each of the two instruments with $k = 1, 2$, both instruments take the form

$$\Delta z^d_{it,(k)} = \Delta \ln\Big(\sum_{j\in J_d} \Omega_{ij,(k)} E_{jt}\Big(\frac{1}{\xi_{ijt}}\frac{\tau_{ij}}{P_{jt}}\Big)^{-(\sigma-1)}\Big), \tag{E.16}$$

where J_d denotes the set of all domestic markets, and $\Omega_{ij,(k)}$ is the weight assigned to market j for firm i according to instrument k.

For the instrument $\Delta z^d_{it,(1)}$, the weights are

$$\Omega_{ij,(1)} \equiv \tau_{ij}^{\widehat{\alpha}_1}, \qquad \text{for all } i \text{ and } j \tag{E.17}$$

where, as a reminder, τ_{ij} captures the trade costs from the municipality where i is located to market j, and $\widehat{\alpha}_1$ is the OLS estimate of α_1 in the estimating equation

$$\ln(R_{ij1}) = \delta_{j,1} + \alpha_1 \ln(\tau_{ij}) + \varepsilon_{ij1}, \qquad \mathbb{E}[\varepsilon_{ij1}|\tau_{ij}] = 0, \tag{E.18}$$

where $\delta_{j,1}$ denotes the destination-j and year-1 fixed effect. When estimating the parameter α_1, we use only information for those firms and destinations such that $j \in J_d$ and $R_{ij1} > 0$.

For instrument $\Delta z^d_{it,(2)}$, the weights are

$$\Omega_{ij,(2)} \equiv R_{ij1}/R_{id1}, \tag{E.19}$$

or, equivalently, the share of the aggregate domestic sales of firm i at period 1 that correspond to the domestic market j.

In words, the two instrumental variables defined in equations (E.16) to (E.19) correspond to the log difference between periods $t = 0$ and $t = 1$ in a weighted sum of firm- and market-specific demand shifters. These two instruments differ only in the weights. The first one assigns a weight to each municipality j that is a function of the trade costs between the municipality of location of firm i and municipality j. The second one assigns a weight to each municipality j that is a function of the share of aggregate domestic revenue of firm i obtained in municipality j at period $t = 1$.

The two instrumental variables defined in equations (E.16) to (E.19) are very similar to those we use in our empirical application. Specifically, the instrument defined by equations (E.16), (E.17) and (E.18) is very similar to both our baseline instrument as well as to those instruments used to compute the estimates reported in columns 2 to 4 of Table 7. The instrument defined in equations (E.16) and (E.19) corresponds to that used in column 5 of Table 7. The reason we use period $t = 1$ information in the construction of the weights $\Omega_{ij,(k)}$ for both instrumental variables $k = 1$ and $k = 2$ is that, in our empirical application, we only observe data on firm-specific sales across domestic markets (municipalities) for one year, and this year is within our sample period.

We expect both the OLS and the two IV estimators defined in equations (E.16) to (E.19) to be biased. The source of this bias is the dependency of the error term in the structural equation, denoted as $\Delta\nu^x_{it}$ in equation (E.15), on the optimal set of export destinations of firm i in period t, J_{ixt}. As the marginal cost function in equation (E.3) is non-constant (i.e $\lambda \neq 0$), any variable that affects a firm's sales in any of the domestic markets will affect the marginal cost at which this firm can sell in any foreign market and, thus, will affect the set of markets to which this firm decides to export. Thus, it is not possible to find an instrumental variable that is both relevant and valid. While this is true, we illustrate in Appendix E.2.3 that, for the two instrumental variables defined in equations (E.16) to (E.19), the bias in the estimate of $-\lambda(\sigma - 1)/(1 + \lambda)$ is small for most values of the structural parameters.

E.2.3 Properties of Estimators: Simulation Results

For each parameterization we consider, we simulate our model 500 times. For each simulation, we solve the model for two periods, $t = 0$ and $t = 1$, and 8,000 firms. In terms of the number of domestic and foreign markets, we consider the following configurations: (a) five domestic markets and five export markets; (b) seven domestic markets and seven export markets. Extending the set of markets beyond fourteen is computationally demanding, as determining the optimal set of domestic and export markets in which a firm i sells in a period t requires solving a combinatorial discrete choice problem with substitutabilities and multiple sources of heterogeneity. Given that the profit function features decreasing differences in the extensive margin of exports, in principle we could have implemented the iterative algorithm in Arkolakis and Eckert (2017) to scale up our exercise, but our simulations results are not much affected by moving from 10 to 14 countries, so we have not pursued this approach here.

We set the elasticity of substitution to equal six (i.e., $\sigma = 6$) and the parameter determining the slope of the marginal cost function with respect to total output to equal 0.904 (i.e., $\lambda = 0.904$). This implies that the elasticity of exports with respect to demand-driven changes in total sales is equal to $-\lambda(\sigma-1)/(1+\lambda) = -2.374$, which coincides with the point estimate reported in column 3 of Table 10.

We also impose the following distributional assumptions:

$$\ln\left(E_{jt}(P_{jt})^{\sigma-1}\right) = 0, \tag{E.20a}$$

$$\ln(\omega_{it}/\varphi_{it}) \sim \mathbb{N}(0,1), \tag{E.20b}$$

$$\ln\left((\tau_{ij})^{-(\sigma-1)}\right) \sim \mathbb{N}(0,1), \tag{E.20c}$$

$$\ln\left((\xi_{ijt})^{\sigma-1}\right) \sim \mathbb{N}(0,\sigma_\xi^2), \qquad \text{with } \sigma_\xi^2 = \{1,5\}, \tag{E.20d}$$

with $\ln(\omega_{it}/\varphi_{it})$ independent across firms and years, $\ln\left((\tau_{ij})^{-(\sigma-1)}\right)$ independent across firms and markets, and $\ln\left((\xi_{ijt})^{\sigma-1}\right)$ independent across firms, markets and years. In equation (E.20a), we eliminate all randomness in the country- and year-specific shifter $E_{jt}(P_{jt})^{\sigma-1}$ because, as the number of countries we can accommodate in our simulation is very small, the results would be entirely driven by the few random draws of this variable. In equation (E.20d), we denote the variance of $\ln\left((\xi_{ijt})^{\sigma-1}\right)$ as σ_ξ^2; we present simulation results for two different values of the variance parameter σ_ξ^2, $\sigma_\xi^2 = 1$ and $\sigma_\xi^2 = 5$.

Concerning the fixed costs to firm i of selling in market j at period t, we impose that one of the domestic markets has zero fixed costs and, for the remaining domestic markets, we impose the following distributional assumption

$$F_{ijt} = f_{ijt}^{\frac{(\sigma-1)(1+\lambda)}{1+\lambda\sigma}} \qquad \text{with} \qquad f_{ijt} \sim \mathbb{N}(0,\sigma_f^2) \quad \text{and} \quad \sigma_f^2 = \{1,5\}, \tag{E.21}$$

with $\tilde{f}_{ijt}$ is independent across firms, markets, and years. We present simulation results for two different values of the variance parameter σ_f^2, $\sigma_f^2 = 1$ and $\sigma_f^2 = 5$.

Concerning the fixed costs of selling in export markets, we consider two different sets of assumptions. First, a case in which we treat foreign markets analogously to domestic markets; i.e., we impose that one foreign market has zero fixed costs and, for the remaining export markets, we impose the distributional assumption in equation (E.21). Second, a case in which we assume that fixed costs for all foreign markets follow the distributional assumption in equation (E.21); i.e., we do not restrict the fixed costs of any foreign market to equal zero. These two models differ in that only in the former will it be true that all firms have positive aggregate exports in every period t; i.e., $R_{ixt} > 0$ for all i and t. Conversely, when fixed costs do not equal zero for all foreign markets, there are some firms that decide to not sell in any foreign market in some period t; for these firms, it is the case that $R_{ixt} = 0$ and, consequently, they are not used in the estimation of $-\lambda(\sigma-1)/(1+\lambda)$.[5]

In Table E.1, we present simulation results for 12 different versions of our model. Each version differs on whether fixed costs for one foreign market are set to zero (indicated in column 1), the value of σ^2_ξ (in column 2); the value of σ^2_f (in column 3); and the total number of markets (in column 4). In column 5, we indicate the average, median and standard deviation of the OLS estimates of the parameter $-\lambda(\sigma-1)/(1+\lambda)$ in the estimating equation in equation (E.15). In column 6 and column 7, we present the same summary statistics for two instrumental variable estimates of $-\lambda(\sigma-1)/(1+\lambda)$; the IV estimate whose distribution is described in column 6 uses as instrument the variable defined in equations (E.16) and (E.17), the IV estimate whose distribution is described in column 7 uses as instrument the variable defined in equations (E.16) and (E.19).

We can extract several lessons from the results in Table E.1. First, the OLS estimator of the parameter of interest $-\lambda(\sigma-1)/(1+\lambda)$ is always biased positively; in fact, although the true value of this parameter is -2.374, the OLS point estimate is always positive and the standard deviation of these OLS point estimates across the 500 simulations is relatively small. Second, for most data generating process considered in our analysis, both IV estimators yield similarly distributed point estimates; specifically, in both cases, the average and the median IV point estimates tend to be smaller than the true parameter value. Third, in models in which the value of σ^2_f is large relative to the value of σ^2_ξ, the downward bias affecting the two IV estimators considered in Table E.1 is very small. Fourth, in models in which only a subset of all active firms export (i.e., models in which the fixed costs of selling to every foreign market follow the distribution in equation (E.21)), the downward bias affecting the two IV estimators can be quantitatively important. Fifth, for all parameterizations we consider, the distribution of the OLS estimator as well as the distributions of the two IV estimators do not vary substantially as we change the number of countries J.

E.3 Biases in the Extensive Margin of Exports

We extend here the analysis in section I to the study of the effect of domestic demand shocks on the extensive margin of exports.

Given the CES demand function in equation (1) and the assumption that firms are monopolistically competitive in every market, firm i will find it profitable to export at time t only if export revenue R_{ixt} exceeds a multiple σ of the fixed cost of exporting F_{ixt}. We can thus express a dummy taking value one if firm i exports at period t as $d_{ixt} = \mathbb{1}\{\ln R_{ixt} > \sigma \ln F_{ixt}\}$, where $\mathbb{1}\{A\}$ denotes an indicator function that takes value one if and only if the statement A is true. The probability that firm i exports conditional on a vector $X_{ix} \equiv \{X_{ixt}\}_t$ that includes a set of period- and sector-specific fixed effects and observed proxies φ^*_{it} and ω^*_{it} for every period t is

$$\Pr(d_{ixt} = 1|X_{ix}) = \mathbb{E}[d_{ixt}|X_{ix}] = \mathbb{E}[\mathbb{1}\{\ln R_{ixt} > \sigma \ln F_{ixt}\}|X_{ix}].$$

Focusing on a linear probability model, we further rewrite the probability of firm i exporting at period t as

$$\Pr(d_{ixt} = 1|X_{ix}) = \mathbb{E}[\ln R_{ixt} - \sigma \ln F_{ixt}|X_{ix}].$$

Therefore, we can write the change in the probability of exporting between any two periods t and $t-1$ as a function of the changes in the log export revenues and log fixed export costs

$$\Pr(d_{ixt} = 1|X_{ix}) - \Pr(d_{ixt} = 0|X_{ix}) = \Delta \Pr(X_{ix}) = \mathbb{E}[\Delta \ln R_{ix} - \sigma \Delta \ln F_{ix}|X_{ix}]$$

[5]For these firms, the dependent variable of interest $\Delta \ln(R_{ixt})$ is not well-defined.

Table E.1: Simulation Results

Model version	One foreign market with $F_{ijt} = 0$?	σ_ξ^2	σ_f^2	J	Summary statistic	$\hat{\beta}_{ols}$	$\hat{\beta}_{iv,1}$	$\hat{\beta}_{iv,2}$
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					Avg.	0.337	-2.643	-3.053
1	Yes	1	1	10	*Med.*	0.336	-2.641	-3.060
					Std. dev.	0.048	0.153	0.136
					Avg.	0.500	-3.073	-2.684
2	No	1	1	10	*Med.*	0.497	-3.056	-2.662
					Std. dev.	0.098	0.368	0.374
					Avg.	0.657	-2.391	-2.485
3	Yes	1	5	10	*Med.*	0.657	-2.391	-2.486
					Std. dev.	0.046	0.137	0.137
					Avg.	0.827	-2.422	-2.501
4	No	1	5	10	*Med.*	0.827	-2.417	-2.501
					Std. dev.	0.056	0.177	0.167
					Avg.	0.386	-2.465	-3.163
5	Yes	5	1	10	*Med.*	0.389	-2.465	-3.168
					Std. dev.	0.045	0.097	0.144
					Avg.	0.338	-2.984	-3.024
6	No	5	1	10	*Med.*	0.338	-2.986	-3.025
					Std. dev.	0.064	0.138	0.186
					Avg.	0.328	-2.792	-3.293
7	Yes	1	1	14	*Med.*	0.329	-2.782	-3.303
					Std. dev.	0.046	0.164	0.165
					Avg.	0.514	-3.232	-2.985
8	No	1	1	14	*Med.*	0.514	-3.242	-2.968
					Std. dev.	0.088	0.348	0.836
					Avg.	0.697	-2.367	-2.461
9	Yes	1	5	14	*Med.*	0.695	-2.374	-2.469
					Std. dev.	0.048	0.152	0.128
					Avg.	0.880	-2.398	-2.498
10	No	1	5	14	*Med.*	0.918	-2.396	-2.495
					Std. dev.	0.055	0.162	0.176
					Avg.	0.426	-2.599	-3.311
11	Yes	5	1	14	*Med.*	0.424	-2.597	-3.315
					Std. dev.	0.052	0.099	0.147
					Avg.	0.376	-2.968	-3.156
12	No	5	1	14	*Med.*	0.376	-2.968	-3.166
					Std. dev.	0.061	0.132	0.167

Note: For each of the models indicated in column 1, results are based on 500 simulations. *Avg.*, *Med.* and *Std. dev.* denote the average, median, and standard deviation, respectively, of the different estimates of $-\lambda(\sigma-1)/(1+\lambda)$. $\hat{\beta}_{ols}$ denotes the OLS estimate; $\hat{\beta}_{iv,1}$ denotes the IV estimate that uses the expression defined in equations (E.16) and (E.17) as instrument; $\hat{\beta}_{iv,2}$ denotes the IV estimate that uses the expression defined in equations (E.16) and (E.19) as instrument. The true value of the parameter is $-\lambda(\sigma-1)/(1+\lambda) = -2.374$. In column 2, we indicate whether fixed costs are set to zero for one of the foreign markets. In column 3, we indicate the value of the parameter σ_ξ^2; in column 4, we indicate the value of the parameter σ_f^2; in column 5, we indicate the total number of markets (both domestic and foreign).

where, from equation (6),

$$\Delta \ln R_{ixt} = \gamma_{sx} + (\sigma - 1)\, \delta_\varphi \Delta \ln(\varphi^*_{it}) - (\sigma - 1)\, \delta_\omega \Delta \ln(\omega^*_{it}) + \varepsilon_{ix},$$

with the different terms in this expression defined as in section I and, analogously

$$\Delta \ln F_{ixt} = \phi_{sx} + \phi_\varphi \Delta \ln \varphi^*_i + \phi_\omega \Delta \ln(\omega^*_i) + u^F_i.$$

Notice that we are being quite flexible, letting firm-level fixed export costs depend on firm-level productivity and factor costs, and on sector fixed effects.

With these expressions at hand, we can write the change in the probability of exporting, expanded to include log domestic sales as an additional covariate, as

$$\begin{aligned} \Delta \Pr(X_{ix}) = \mathbb{E}\big[&(\gamma_{sx} - \phi_{sx}) + (\gamma_{\ell x} - \phi_{\ell x}) + [(\sigma - 1)\, \gamma_\varphi - \sigma\phi_\varphi]\Delta \ln(\varphi^*_i) \\ &- [(\sigma - 1)\, \delta_\omega - \sigma\phi_\omega]\Delta \ln(\omega^*_i) + \beta \Delta \ln R_{id} + \varepsilon_{ix} - \sigma u^F_i \big| X_{ix}\big], \end{aligned}$$

where, as in equation (9), $\varepsilon_{ix} = (\sigma - 1)\, [u^\xi_{ix} + u^\varphi_i - u^\omega_i]$. Following the same steps as in section I, the following asymptotic properties of $\hat{\beta}_{OLS}$ can be derived:

$$plim(\hat{\beta}_{OLS}) = \frac{cov(u^\xi_{ix} + u^\varphi_i - u^\omega_i - \frac{\sigma}{\sigma-1} u^F_i, u^\xi_{id} + u^\varphi_i - u^\omega_i)}{var(u^\xi_{id} + u^\varphi_i - u^\omega_i)}.$$

The only difference relative to equation (11) is the addition of the term $-(\sigma/(\sigma - 1))u^F_i$ in the first element of the covariance in the numerator. It is clear that, as in the intensive margin regressions, this covariance is likely to be positive, thus generating a positive value of $plim(\hat{\beta}_{OLS})$.

The probability limit of the IV estimator of β is given by

$$plim(\hat{\beta}_{IV}) = \frac{cov(u^\xi_{ix} + u^\varphi_i - u^\omega_i - \frac{\sigma}{\sigma-1} u^F_i, \mathcal{Z}_{id})}{cov(u^\xi_{id} + u^\varphi_i - u^\omega_i, \mathcal{Z}_{id})}. \tag{E.22}$$

This expression will equal zero as long as the instrument $\mathcal{Z}_{id}$ verifies the following two conditions: (a) it is correlated with the boom-to-bust change in domestic sales of firm i, after partialling out firm fixed effects and the boom-to-bust difference in observable determinants of the firm's marginal cost; and (b) it is mean independent of the boom-to-bust changes in unobserved productivity, u^φ_i, factor costs, u^ω_i, export demand shocks, u^ξ_{ix}, and export fixed-cost shocks u^F_i (this latter being the only additional condition relative to our results for the intensive margin regressions). As in our discussion in section I, an instrument can only (generically) verify conditions (a) and (b) if its effect on domestic sales works exclusively through the domestic demand shock u^ξ_{id}.

It is straightforward to extend the above analysis to the case in which total sales and exports (but not the dummy variable d_{ixt} indicating whether firm i exports at period t) are measured with error and domestic sales are imputed by subtracting exports from total sales. Following the same steps as in Appendix E.1, we obtain

$$plim(\hat{\beta}_{IV}) = \frac{cov(u^\xi_{ix} + u^\varphi_i - u^\omega_i - \frac{\sigma}{\sigma-1} u^F_i + \frac{1}{\sigma-1}\varpi_{ix}, \mathcal{Z}_{id})}{cov(u^\xi_{id} + u^\varphi_i - u^\omega_i + \frac{1}{\sigma-1}\left(\varpi_{iT} - \varpi_{ix}\right), \mathcal{Z}_{id})}.$$

Given that the numerator in this expression coincides with that in equation (E.22), the presence of measurement error in total sales and exports does not affect the conditions that the instrument Z_{id} must satisfy so that the probability limit of the IV estimator equals zero. Thus, as long as the conditions (a) and (b) above are satisfied, $plim(\hat{\beta}_{IV}) = 0$ independently of the relationship between the instrument and the measurement errors in total sales and exports.

E.4 The Relevance of Confounding Export Demand Shocks

As formalized in section I.A (see discussion of equation (12)) and in section VI.A, a possible source of bias affecting our TSLS estimates of the elasticity of changes in firms' exports with respect to changes in their domestic (or total) sales is the possible non-zero correlation between our instrument and the changes foreign demand affecting each firm. Testing whether this non-zero correlation is present in our empirical setting is complicated by the fact that firm-specific export-demand shocks are not directly observed in our data.

Different firms may face different foreign demand shocks for two reasons: (a) they produce different goods; (b) they sell in different foreign countries. To the extent that good-specific demand shocks are adequately controlled for by the sector-specific fixed effects we include in all our regression specifications, heterogeneity in foreign demand shocks due to reason (a) will not bias our TSLS estimates of the elasticity of firms' exports with respect to their domestic (or total) sales. Here, we study whether our estimates may be biased due to the heterogeneity in the set of foreign countries that firms export to.

Specifically, we explore here whether the boom-to-bust changes in the number of vehicles per capita in a municipality is correlated with a municipality-specific aggregator of the boom-to-bust demand shocks experienced by different foreign countries.

To construct municipality- and period-specific measures of foreign demand shocks, we follow a four-step procedure. First, for each destination country to which Spain exported a positive amount in 2002, we collect 2002-2013 data from UN Comtrade (United Nations, n.d.) on its aggregate imports by product, country of origin and year, with a "product" corresponding either to an HS-2, an HS-4 or an HS-6 digit product code. Second, we regress the logarithm of this product-, origin-, destination-, and year-specific import measure on origin- and year-specific fixed effects, destination- and year-specific fixed effects and product- and year-specific fixed effects. We interpret the estimates of the destination- and year-specific fixed effects as estimates of destination- and year-specific demand shifters after controlling for sectoral shifters. Third, we compute municipality-specific weighted averages of these destination- and year-specific demand shifters, where the weight that each destination country takes for each municipality equals the share of the 2002 exports of that municipality to that destination country. Fourth, we compute municipality-specific measures for the boom and bust periods as the average of the 2002-2008 years and the 2009-2013 years, respectively.

In Table E.2, we present OLS estimates of the coefficients in regressions of the boom-to-bust log change in the measure of foreign demand shocks whose construction is described in the previous paragraph on the boom-to-bust log change in the number of vehicles per capita. Each observation in these regressions corresponds to a municipality, and we weight each municipality by the number of firms located in the corresponding municipality that have positive exports in the boom and bust periods. The results show that, even if we use a 10% significance level test, we cannot reject the null hypothesis that the boom-to-bust log change in our measure of municipality-specific foreign demand shocks is uncorrelated with the boom-to-bust log change in the number of vehicles per capita in the corresponding municipality.

Table E.2: Correlation of Local and Foreign Demand Shocks

Product Definition:	HS-2 (1)	HS-4 (2)	HS-6 (3)
Δ Ln(Vehicles p.c.)	-2.842	-0.622	-0.188
	(2.180)	(0.469)	(0.177)
	(2.261)	(0.474)	(0.175)
Obs.	1,103	1,103	1,103

Notes: [a] denotes significance at the 1% level; [b] denotes significance at the 5% level; [c] denotes significance at the 10% level. In parenthesis, we report standard errors. The first set of standard errors are heteroskedasticity-robust standard errors; the second set are standard errors clustered by province.

F Estimation of Revenue Productivity

We present a step-by-step description of our baseline estimation approach in Appendix F.1. For an analogous description of the alternative estimation approach used to compute the estimates in columns 3 and 4 of Table 9, see Bilir and Morales (2020). We summarize the production function estimates that both approaches yield in Appendix F.2.

F.1 Baseline Estimation Approach

We describe here the procedure we follow to estimate a proxy for firm- and year-specific performance or revenue productivity under the assumption that the production function is Leontief in materials. We describe first the assumptions that we impose on the production function, the demand function, market structure, and the stochastic process of revenue productivity or performance. Given these assumptions, we illustrate how we estimate the demand elasticity σ and all parameters of the revenue function. Finally, we describe how we use these estimates to recover a proxy of the revenue productivity or performance for every firm and year.

Assumption on production function. We assume a production function that is a Leontief function of materials and a translog aggregator of labor and capital (as in Ackerberg, Caves and Frazer, 2015):

$$Q_{it} = \min\{H(K_{it}, L_{it}; \boldsymbol{\alpha}), M_{it})\}\varphi_{it}, \tag{F.1a}$$

$$H(K_{it}, L_{it}; \boldsymbol{\alpha}) = \exp(h(k_{it}, l_{it}; \boldsymbol{\alpha})), \tag{F.1b}$$

$$h(k_{it}, l_{it}; \boldsymbol{\alpha}) \equiv \alpha_l l_{it} + \alpha_k k_{it} + \alpha_{ll} l_{it}^2 + \alpha_{kk} k_{it}^2 + \alpha_{lk} l_{it} k_{it}, \tag{F.1c}$$

with $\boldsymbol{\alpha} = (\alpha_l, \alpha_k, \alpha_{ll}, \alpha_{kk}, \alpha_{lk})$. In equation (F.1a), K_{it} is effective units of capital, L_{it} is the number of production workers, M_{it} is a quantity index of materials use, and φ_{it} denotes the Hicks-neutral physical productivity. To simplify the notation, we use here lower-case Latin letters to denote the logarithm of the upper-case variable, e.g., $l_{it} = \ln(L_{it})$. The production function in equation (F.1) nests that introduced in Appendix A, which implicitly assumes that $\alpha_{ll} = \alpha_{kk} = \alpha_{lk} = 0$. In our estimation, we impose no *a priori* restriction on the values of the elements of the parameter vector $\boldsymbol{\alpha}$ and, thus, our estimation framework does not take a stand on whether marginal production costs are constant (as assumed in section I) or increasing (as assumed in section VI).

Consistently with the definition of φ_{it} as physical productivity, we assume that

$$\mathbb{E}[\varphi_{it}|\mathcal{J}_{it}] = \varphi_{it}, \tag{F.2}$$

where $\mathcal{J}_{it}$ denotes the information set of firm i at the time at which the period-t pricing and input decisions are taken. Therefore, the firm knows the value of its productivity φ_{it} when making the period-t pricing and input decisions.

We assume that both materials and labor are fully flexible inputs, and that capital is dynamic and determined one period ahead. Consequently, both M_{it} and L_{it} are a function of $\mathcal{J}_{it}$, while K_{it} is a function of $\mathcal{J}_{it-1}$.

Assumptions on demand function. We assume that firms face a constant elasticity of substitution demand function as described in equation (1), and impose the assumption that the demand shock ξ_{it} is known to firms when determining their input and output decisions; i.e.,

$$\mathbb{E}[\xi_{it}|\mathcal{J}_{it}] = \xi_{it}. \tag{F.3}$$

Assumptions on market structure. As described in section (2), we assume that firms are monopolistically competitive in the output markets and that they take the prices of labor, materials and capital as given.

Derivation of the revenue function. Given the assumption that materials is a flexible input, equation (F.1a) implies that optimal materials usage satisfies

$$M_{it} = H(K_{it}, L_{it}; \boldsymbol{\alpha}).$$

Therefore, we can rewrite the production function in equation (F.1a) as

$$Q_{it} = H(K_{it}, L_{it}; \boldsymbol{\alpha})\varphi_{it}, \tag{F.4}$$

where $H(K_{it}, L_{it}; \boldsymbol{\alpha})$ is defined as in equations (F.1b) and (F.1c). Given this expression and the demand function in equation (1), we can write the revenue function of a firm i at period t as

$$R_{it} = P_{it}Q_{it} = P_{st}^{\frac{\sigma-1}{\sigma}} E_{st}^{\frac{1}{\sigma}} \xi_{it}^{\frac{\sigma-1}{\sigma}} Q_{it}^{\frac{\sigma-1}{\sigma}} = \mu_{st} H(K_{it}, L_{it}; \boldsymbol{\beta})\psi_{it}, \tag{F.5}$$

where

$$\kappa \equiv (\sigma - 1)/\sigma, \tag{F.6a}$$

$$\boldsymbol{\beta} \equiv \kappa\boldsymbol{\alpha}, \tag{F.6b}$$

$$\psi_{it} \equiv (\xi_{it}\varphi_{it})^{\kappa} \tag{F.6c}$$

$$\mu_{st} \equiv P_{st}^{\kappa}(E_{st})^{1-\kappa}. \tag{F.6d}$$

The parameter κ measures the inverse of the firm's markup. While the parameter vector $\boldsymbol{\alpha}$ includes the production function parameters, the vector $\boldsymbol{\beta}$ includes the revenue function parameters. The variable ψ_{it} captures the revenue productivity of the firm: the residual determinant of a firm's revenue after controlling for sector- and year-specific fixed effects and for the effect of capital and labor on the firm's revenue. As illustrated in equation (F.6c), revenue productivity equals in our model the product of the Hicks-neutral productivity φ_{it} and the demand shifter ξ_{it} to the power of the reciprocal of the firm's markup. The sector-year fixed effects accounts for the price index and total expenditure in the corresponding sector-year pair.

Assumptions on stochastic process for revenue productivity. We assume that revenue productivity follows a first-order autoregressive process, AR(1), with a state- and year-specific shifter:

$$\psi_{it} = \gamma_{st} + \rho\psi_{it} + \eta_{it} \qquad \text{with} \qquad \mathbb{E}[\eta_{it}|\mathcal{J}_{it}] = 0. \tag{F.7}$$

This stochastic process for revenue productivity may arise under different stochastic process for physical productivity φ_{it} and the demand shifter ξ_{it}; e.g., both variables follow AR(1) process with identical persistence parameters equal to ρ; or, one of them follows an AR(1) process with persistence parameter ρ and the other one is independent over time.

Estimation of demand elasticity. In order to estimate the demand elasticity σ, we follow the approach implemented, among others, in Das, Roberts and Tybout (2007) and Antras, Fort and Tintelnot (2017). Given the assumption that all firms are monopolistically competitive in their output markets, it will be true that

$$R_{it} - C_{it}^{v} = \frac{1}{\sigma} R_{it},$$

where C_{it}^{v} denotes the total variable costs that firm i incurred at period t to obtain the sales revenue R_{it}. This expression indicates that the firm's total profits (gross of fixed costs) is equal to the reciprocal of the demand elasticity of substitution σ multiplied by the firm's total revenues. Given that the only variable inputs are materials M_{it} and labor L_{it}, we can rewrite this relationship as

$$R_{it} - P_{it}^{m} M_{it} - \omega_{it} L_{it} = \frac{1}{\sigma} R_{it},$$

where P_{it}^{m} denotes the equilibrium materials' price faced by firm i at period t, ω_{it} denotes the equilibrium salary and, thus, $P_{it}^{m} M_{it}$ denotes total expenditure in materials' purchases and $\omega_{it} L_{it}$ denotes total payments to labor. Rearranging terms, we obtain the following equality

$$\left(\frac{\sigma - 1}{\sigma}\right) R_{it} = P_{it}^{m} M_{it} + \omega_{it} L_{it},$$

and, allowing for measurement error in sales revenue, $R_{it}^{obs} \equiv R_{it} \exp(\varepsilon_{it})$, we obtain

$$\ln\left(\frac{\sigma - 1}{\sigma}\right) + r_{it}^{obs} - \varepsilon_{it} = \ln(P_{it}^{m} M_{it} + \omega_{it} L_{it}),$$

where, as indicated above, lower-case Latin letters denote the logarithm of the corresponding upper case variable and, thus, $r_{it}^{obs} \equiv \ln(R_{it}^{obs})$. Imposing the assumption that $\mathbb{E}[\varepsilon_{it}] = 0$, we identify σ through the following moment condition

$$\mathbb{E}\Big[\ln\left(\frac{\sigma - 1}{\sigma}\right) + r_{it}^{obs} - \ln(P_{it}^{m} M_{it} + \omega_{it} L_{it})\Big] = 0. \tag{F.8}$$

Estimation of labor elasticity parameters. Given equation (F.5), we can write the profit function of firm i in period t as

$$\Pi_{it} = \mu_{st} H(K_{it}, L_{it}; \boldsymbol{\beta}) \psi_{it} - \omega_{it} L_{it} - P_{it}^{m} M_{it} - P_{it}^{k} I_{it},$$

where ω_{it} denotes the wage that firm i faces at period t and, analogously, P_{it}^{m} and P_{it}^{k} denote the materials and capital prices. Assuming that labor is a fully flexible input and that firms are both monopolistically competitive in output markets and take the price of all inputs as given, the first order condition of the profit function with respect to labor implies that

$$\frac{\partial \Pi_{it}}{\partial L_{it}} = (\beta_l + 2\beta_{ll} l_{it} + \beta_{lk} k_{it}) R_{it} - \omega_{it} L_{it} = 0.$$

Reordering terms and taking logs on both sides of the equality, we obtain

$$\ln(\beta_l + 2\beta_{ll} l_{it} + \beta_{lk} k_{it}) = \ln(\omega_{it} L_{it}) - r_{it},$$

and, taking into account that revenues are measured with error, we can further rewrite

$$\ln(\beta_l + 2\beta_{ll} l_{it} + \beta_{lk} k_{it}) = \ln(\omega_{it} L_{it}) - r_{it}^{obs} + \varepsilon_{it}.$$

Assuming that the measurement error in revenue is not only mean zero (as imposed to derive the moment condition in equation (F.8)) but mean independent of the firm's labor and capital usage,

$$\mathbb{E}[\varepsilon_{it} | l_{it}, k_{it}] = 0,$$

we can derive the following conditional moment:

$$\mathbb{E}[r_{it}^{obs} - \ln(\omega_{it} L_{it}) + \ln(\beta_l + 2\beta_{ll} l_{it} + \beta_{lk} k_{it}) | l_{it}, k_{it}] = 0.$$

We derive unconditional moments from this equation and use a method of moments estimator to estimate $(\beta_l, \beta_{ll}, \beta_{lk})$. With the estimates $(\hat{\beta}_l, \hat{\beta}_{ll}, \hat{\beta}_{lk})$ in hand, we recover an estimate of the measurement error ε_{it} for each firm i, affiliate j, and period t:

$$\hat{\varepsilon}_{it} = r_{it}^{obs} - \ln(\omega_{it} L_{it}) + \log(\hat{\beta}_l + 2\hat{\beta}_{ll} l_{it} + \hat{\beta}_{lk} k_{it}).$$

Combining the estimates of the parameters entering the elasticity of the firm's revenues with respect to labor, $(\hat{\beta}_l, \hat{\beta}_{ll}, \hat{\beta}_{lk})$, and the estimate of the demand elasticity of substitution, we compute estimates of the parameters $(\alpha_l, \alpha_{ll}, \alpha_{lk})$; i.e.,

$$(\hat{\alpha}_l, \hat{\alpha}_{ll}, \hat{\alpha}_{lk}) = \frac{\hat{\sigma}}{\hat{\sigma} - 1} (\hat{\beta}_l, \hat{\beta}_{ll}, \hat{\beta}_{lk}).$$

Estimation of capital elasticity parameters. Using the estimates $(\hat{\beta}_l, \hat{\beta}_{ll}, \hat{\beta}_{lk})$ and $\hat{\varepsilon}_{it}$ we can construct a corrected measure of revenues

$$\widehat{r}_{it} \equiv r_{it} - \hat{\beta}_l l_{it} - \hat{\beta}_{ll} l_{it}^2 - \hat{\beta}_{lk} l_{it} k_{it} - \hat{\varepsilon}_{it},$$

and, given the expression for sales revenues in equation (F.5), it holds that

$$\widehat{r}_{it} = \beta_k k_{it} + \beta_{kk} k_{it}^2 + \psi_{it}.$$

Given this expression and the stochastic process for the evolution of productivity in equation (F.7), it will be true that

$$\widehat{r}_{it} = \beta_k k_{it} + \beta_{kk} k_{it}^2 + \mu_\psi (\widehat{r}_{ijt-1} - \beta_k k_{ijt-1} - \beta_{kk} k_{ijt-1}^2) + \zeta_{st} + \eta_{it}, \tag{F.9}$$

where ζ_{st} is an unobserved sector- and time-specific effect that accounts for the revenue shifter μ_{st} and the productivity shifter γ_{st}. Given that both L_{it} and K_{it} are a function of the information set $\mathcal{J}_{it}$, the definition of η_{it} in equation (F.7) implies that

$$\mathbb{E}[\eta_{it} | k_{it}, \widehat{r}_{ijt-1}, \{d_{st}\}_{s,t}] = 0,$$

where $\{d_{st}\}_{s,t}$ denotes a full set of sector- and time-specific dummy variables. Therefore, we can derive the following conditional moment equality

$$\mathbb{E}[\widehat{r}_{it} - \beta_k k_{it} - \beta_{kk} k_{it}^2 - \rho(\widehat{r}_{ijt-1} - \beta_k k_{ijt-1} - \beta_{kk} k_{ijt-1}^2) - \zeta_{st} | k_{it}, \widehat{r}_{ijt-1}, \{d_{st}\}_{s,t}] = 0$$

We derive unconditional moments from this equation and use a method of moments estimator to estimate $(\beta_k, \beta_{kk}, \rho)$. When estimating these parameters, we use the Frisch-Waugh-Lovell theorem to control for the full set of sector- and time-specific fixed effects $\{\zeta_{st}\}_{s,t}$. Combining the estimates of the parameters (β_k, β_{kk}), and the estimate of the demand elasticity of substitution σ, we compute estimates of the parameters (α_k, α_{kk}); i.e.,

$$(\hat{\alpha}_k, \hat{\alpha}_{kk}) = \frac{\hat{\sigma}}{\hat{\sigma} - 1}(\hat{\beta}_k, \hat{\beta}_{kk}).$$

Estimation of productivity. We can also use the estimates of the parameters (β_k, β_{kk}) and the constructed random variable $\widehat{r}_{it}$ to build an estimate of the revenue productivity ψ_{it} for every firm and time period

$$\widehat{\psi}_{it} = \widehat{r}_{it} - \hat{\beta}_k k_{it} - \hat{\beta}_{kk} k_{it}^2.$$

F.2 Production Function Estimates

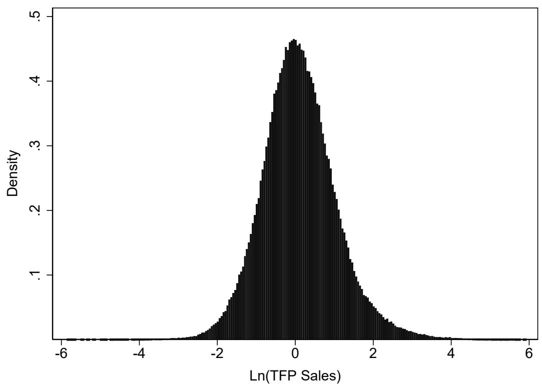
We summarize here the production function and productivity estimates that we obtain both when we assume a production function that is Leontief in materials (see Appendix F.1 for the corresponding estimation approach) and when we assume instead a production function that is Cobb-Douglas in materials (see Bilir and Morales, 2020, for the corresponding estimation approach). No matter which of these two production functions we assume, we estimate the corresponding production function parameters and demand parameters separately for the boom and bust periods and for each of the twenty-four 2-digit NACE sectors in which the manufacturing firms in our dataset are classified. For both the boom and the bust periods, we report here the simple average across all sectors of the estimated labor and capital elasticities, of the estimated persistence parameters ρ and of the demand elasticity σ.

Under the assumption that the production function is Leontief in materials, we obtain the following estimates. In the boom period, the average elasticities of revenue with respect to labor and capital are 0.23 and 0.19, respectively; the average annual autocorrelation in performance is 0.97; and the average demand elasticity is 3.55. In the bust period, the average elasticities of revenue with respect to labor and capital are 0.26 and 0.18, respectively; the average annual autocorrelation in performance is 0.98; and the average demand elasticity is 3.37.

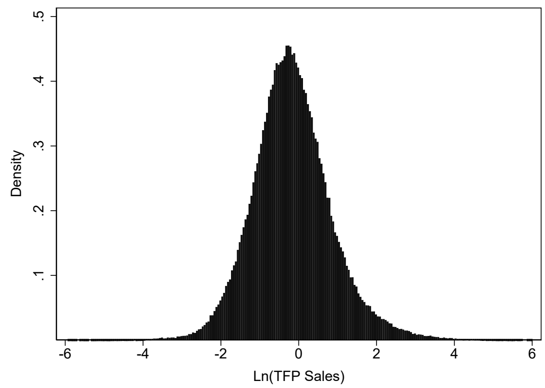
Under the assumption that the production function is Cobb-Douglas in materials, we obtain the following estimates. In the boom period, the average elasticities of value added with respect to labor and capital are 0.76 and 0.19, respectively; the average annual autocorrelation in performance is 0.78; and the average demand elasticity is 3.19. In the bust period, the average elasticities of value added with respect to labor and capital are 0.86 and 0.17, respectively; the average annual autocorrelation in performance is 0.78; and the average demand elasticity is 3.05.

Notice that both estimation approaches yield estimates of the demand elasticity σ that are a bit low relative to those that, using identification strategies different from ours, are typically obtained in the international trade literature (see Head and Mayer, 2014, for a review). One possible explanation for this mismatch between our estimates and those in the international trade literature is the fact that we cannot observe firms' expenditure in energy; this may imply that our measure of the variable production costs underestimates the firms' total expenditure in variable inputs and, thus, that our estimates of σ are downward biased. These estimates of σ do not, however, impact

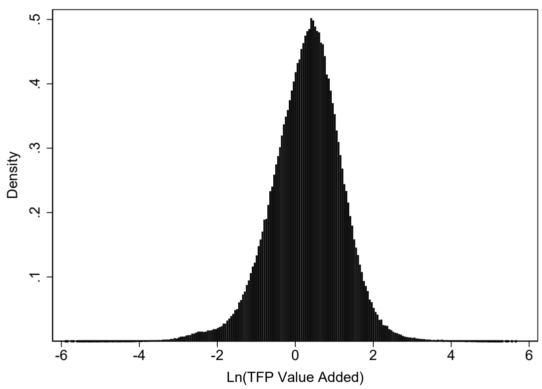
Figure F.1: Productivity Estimates



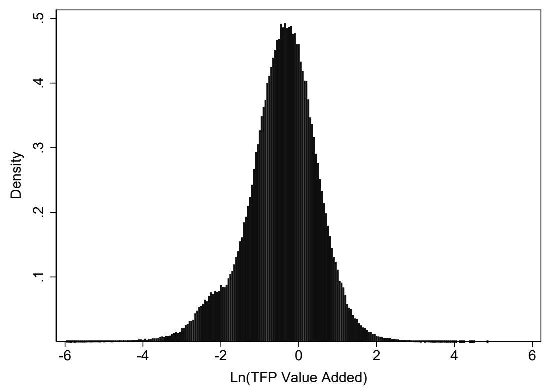
(a) Density of Ln(TFP Sales) in Boom



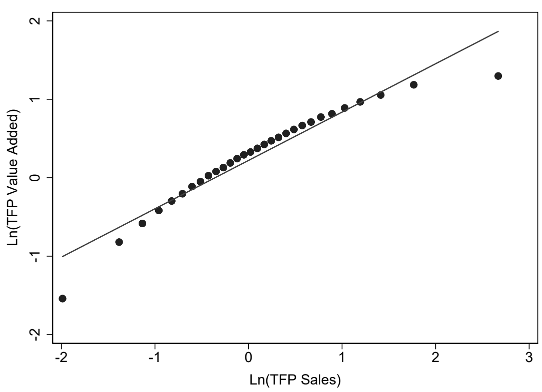
(b) Density of Ln(TFP Sales) in Bust



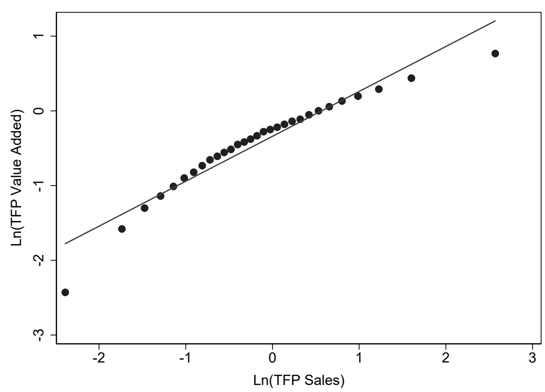
(c) Density of Ln(TFP Value Added) in Boom



(d) Density of Ln(TFP Value Added) in Bust



(e) Binscatter in Boom



(f) Binscatter in Bust

Notes: The figures in panels (a) and (b) present the density function of our (log) TFP estimates in boom and bust, respectively, following the procedure in section F.1.The figures in panels (c) and (d) present the density function of our (log) TFP estimates in boom and bust, respectively, following the procedure in Bilir and Morales (2020). The figure in panel (e) presents a binscatter illustrating the relationship in the boom period between our two estimates of the firm's log TFP. The figure in panel (f) is analogous for the case of the bust period. The slope of the regression lines in panels (e) and (f) are, respectively, 0.62 and 0.6.

any of the estimates presented in the main draft. More specifically, the only exercise that we perform in the paper and that relies on the estimated value of σ is the quantification in section VII. However, as indicated in that section, our baseline quantification calibrates the value of σ to a central value among the estimates computed in the international trade literature; i.e., $\sigma = 5$.

Panels (a) and (b) in Figure F.1 show, respectively, that the marginal distribution in the boom and bust periods of the (log) TFP estimates computed following the procedure in section F.1. These two marginal distributions are symmetric around zero and close to normally distributed, reflecting that the distribution of the TFP estimates is roughly log-normal. Panels (c) and (d) show analogous marginal distributions for (log) TFP estimates computed following the procedure in Bilir and Morales (2020) While the distributions in panels (a) and (b) are similar to each other, that in panel (d) is clearly different from that in panel (c) in that the fraction of firms in the lower tail of the distribution is significantly larger. Thus, our value added-based TFP estimates show that the fraction of firms with relatively lower TFP increased in the bust period relative to the boom.

Panels (e) and (f) in Figure F.1 show how our two measures of TFP relate to each other. They show that there is a positive association between both measures; i.e., firms that have higher TFP according to our sales-based measure also tend to have higher TFP according to our value added measure. However, the relationship between both is not perfectly linear but slightly concave.

G Additional Robustness Tests

G.1 Results with Local (Municipality) Instrument

In this section, we report results corresponding to Tables 5, 8, and 9 in the paper using the local instrument instead of the gravity-based instrument.

Table G.1 reports results from the extensive-margin regressions using the local instrument. The table follows the same structure as Table 5 in the paper. Column 1 reports the first-stage relationship, which is statistically significant and of similar magnitude as with the gravity-based instrument. Columns 2 and 4 report results from OLS specifications, which are identical to those in the corresponding columns of Table 5. Columns 3 and 5 report second-stage coefficients, which are both negative but statistically insignificant. Therefore, the conclusion regarding the extensive margin effect is qualitatively the same as in the main text: the vent-for-surplus mechanism does not appear to operate particularly via the extensive margin (i.e., via entry and exit from the export market).

Table G.1: Extensive Margin: 2SLS Estimates for Local Instrument

Dependent Variable:		Export Dummy		Proportion of Years	
	1st Stage	OLS	2SLS	OLS	2SLS
	(1)	(2)	(3)	(4)	(5)
Ln(Domestic Sales)		0.021^a	-0.212^c	0.010^a	-0.005
		(0.005)	(0.108)	(0.003)	(0.055)
Ln(Vehicles p.c. in municipality)	0.187^a				
	(0.042)				
Ln(TFP)	1.149^a	0.069^a	0.336^a	0.063^a	0.079
	(0.015)	(0.007)	(0.125)	(0.005)	(0.063)
Ln(Av. Wages)	-0.574^a	-0.048^a	-0.181^a	-0.042^a	-0.050
	(0.014)	(0.006)	(0.061)	(0.004)	(0.031)
Observations	126,024	126,024	126,024	126,024	126,024
R-squared	0.982	0.841	-0.092	0.918	0.020
Firm FE	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes
Sector-Period FE	Yes	Yes	Yes	Yes	Yes
F-statistic	20.14				
Mean of Dep. Var		0.173	0.173	0.114	0.114
Ext-Margin Elasticity		0.121	-1.223	0.0848	-0.0413

Note: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered by municipality reported in parenthesis. Ln(Vehicles p.c.) denotes the log of vehicles per capita at the municipal level. F-statistic denotes the corresponding test statistic for the null hypothesis that the coefficient on Ln(Vehicles p.c.) equals zero. All specifications include firm fixed effects, province fixed effects, and sector-period fixed effects. The estimation sample includes all firms selling in the domestic market in at least one year in the period 2002-2008 and in the period 2009-2013.

Table G.2 reports estimates of specifications including confounding factors where we use the Local instrument. The table follows the same structure as Table 8 in the paper. Column 1 reproduces the result from column 8 in panel A of Table 3. In column 2, we include as a potential confounding factor the firm-level change in the share of temporary workers. The negative and significant point estimate is consistent with the one found with the gravity-based Instrument in the paper. In columns 3 and 4, we include municipality-level controls for labor market conditions: the change in the share of temporary workers (column 3) and the change in manufacturing employment per capita (column 4). The inclusion of these controls has little effect on the main coefficient of interest and only the second one is statistically significant, which is again consistent with the results in Table 8.

In columns 5 to 7, we study potential confounding effects related to financial costs. As explained in the paper, our measure of financial costs is the within-period average ratio of financial expenditures over total outstanding debt with financial institutions (both measures are annually reported by firms in their financial statements). In column 5, we add the log change in this firm-level measure of financial costs as an additional control. The point estimate is negative and significant at the 5% level, but the impact on the coefficient of interest is negligible. In columns 6 and 7, we explore the possibility that the relevant increase in the financial costs faced by firms in the bust relative to the boom happened through credit rationing, instead of via explicit interest rates. Regardless of whether we measure financial costs in the boom at the firm level (column 6) or at the municipal level (column 7), our results indicate that either credit rationing had little impact on firms' exports or our conjecture that it may be measured through the firms' financial costs in the boom has little empirical support. Again, these results are qualitatively the same as those reported in Table 8 for the gravity-based instrument.

Table G.2: Confounding Factors: Local IV

Dependent Variable:	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Domestic Sales)	-1.327^a	-1.247^a	-1.417^a	-1.435^a	-1.305^a	-1.303^a	-1.414^a
	(0.337)	(0.342)	(0.381)	(0.377)	(0.373)	(0.362)	(0.379)
ΔShare of Temp. Workers		-0.243^b					
(firm level)		(0.106)					
ΔShare of Temp. Workers			0.028				
(munic. level)			(0.187)				
ΔManufacturing Emp. p.c.				0.224^a			
(munic. level)				(0.051)			
ΔLn(Financial Costs)					-0.031^b		
(firm level)					(0.013)		
Financial Costs in Boom						-0.009	
(firm level)						(0.014)	
Financial Costs in Boom							0.001
(munic. level)							(0.045)
Observations	8,009	7,640	7,743	7,745	6,879	6,945	7,741
F-Statistic	33.10	31.80	29.30	28.30	29.55	29.93	29.50

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by municipality reported in parentheses. In all specifications, ΔLn(Domestic Sales) is instrumented by ΔLn(Vehicles per capita), defined as in previous tables. All specifications include firm-level log changes in TFP and in log wages as additional controls (coefficients not included to save space), and sector and province fixed effects.

Table G.3 reports estimates with two alternative TFP measures. It follows the same structure as Table 9 in the paper, but using the local instrument. Columns 1 and 2 report the results from column 5 of Table 2 and from column 8 in panel A of Table 3 in the paper, where we use a definition of TFP based on total revenue from sales. In columns 3 and 4, we use an alternative definition of TFP based on value added (for details on our TFP measures, see Section V.D in the paper and also F in this online appendix). The OLS coefficient on the change in log domestic sales is close to zero and insignificant, rather than negative as in column 1. The two-stage least squares point estimate is -0.686, which is smaller in magnitude than column 2, but still statistically significant.

Table G.3: Alternative TFP Measures: Local IV

Dependent Variable:		ΔLn(Exports)		
	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
Δ Ln(Domestic Sales)	-0.284^a	-1.327^a	0.027	-0.686^c
	(0.030)	(0.337)	(0.028)	(0.387)
Δ Ln(Average Wages)	-0.712^a	-1.240^a	-0.749^a	-0.925^a
	(0.059)	(0.178)	(0.068)	(0.115)
Δ Ln(TFP)	1.522^a	2.533^a		
	(0.051)	(0.323)		
Δ Ln(TFP Value Added)			1.016^a	1.227^a
			(0.063)	(0.126)
Observations	8,009	8,009	8,009	8,009
F-Statistic		33.10		28.12

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered at the municipality level are reported in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the difference in $\text{Ln}(X)$ between its average in the 2009-2013 period and its average in the 2002-2008 period. All specifications include sector and province fixed effects.

G.2 Panel Results with Lagged Instruments

Table G.4 reports results of yearly regressions that also include lags of the instruments in the first-stage specification. As mentioned in the main text, the instruments (regardless of whether it is the local or the gravity-based one) continue to be weak in this case. A possible explanation for this negative findings is that regions in Spain might differ in the lag structure with which changes in the stock of vehicles per capita correlate with changes in the demand for manufacturing products produced in the corresponding region. As we ignore what the relevant lag is for each specific region, the estimates we obtain for the coefficient on each lag is a combination of the true coefficient of those municipalities for which such lag is relevant and a zero coefficient for those municipalities for which it is not relevant. This could explain why the estimates on the different lags of the stock of vehicles per capita are not significant in these first-stage specifications.

G.3 Alternative Samples

Table G.5 reports results from regressions that estimate our main specification using the gravity-based instrument on a variety of subsamples of continuing exporters. In columns 1 through 3 we report results that, in turn: (i) exclude multinational subsidiaries operating in Spain, (ii) include only firms with a single manufacturing establishment, and (iii) include only firms with a single establishment. In terms of data sources, whether an establishment is part of a multinational firm can be inferred from the fact that the firm-level identifier included in the Commercial Registry

Table G.4: Intensive-Margin Regressions with Lagged Instrument

Panel A: Local Instrument

Instrument:	Including lags of IV				Combined lags of IV			
	1st Stage	2SLS	1st Stage	2SLS	1st Stage	2SLS	1st Stage	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Ln(Domestic Sales)		0.256		0.305		1.517		2.642
		(0.823)		(0.777)		(2.356)		(4.081)
Ln(Vehicles p.c. in municipality)	0.001		-0.053					
	(0.066)		(0.070)					
Ln(Vehicles p.c. in municipality)$_{t-1}$	0.091^{b}		0.104^{c}					
	(0.042)		(0.055)					
Ln(Vehicles p.c. in municipality)$_{t-2}$			0.026					
			(0.062)					
Average of Ln(Vehicles p.c. in municipality) in t and $t-1$					0.042			
					(0.048)			
Average of Ln(Vehicles p.c. in municipality) in $t-1$ and $t-2$							0.036	
							(0.047)	
Ln(TFP)	0.865^{a}	0.933	0.863^{a}	0.909	0.920^{a}	-0.263	0.920^{a}	-1.298
	(0.040)	(0.716)	(0.050)	(0.673)	(0.030)	(2.169)	(0.030)	(3.761)
Ln(Av. Wages)	-0.344^{a}	-0.437	-0.350^{a}	-0.448	-0.373^{a}	0.087	-0.373^{a}	0.506
	(0.035)	(0.290)	(0.044)	(0.278)	(0.029)	(0.879)	(0.029)	(1.528)
Observations	45,384	45,384	35,863	35,863	60,196	60,196	60,196	60,196
F-statistic	2.55		1.95		0.78		0.61	

Panel B: Gravity-based Instrument

Instrument:	Including lags of IV				Combined lags of IV			
	1st Stage	2SLS	1st Stage	2SLS	1st Stage	2SLS	1st Stage	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Ln(Domestic Sales)		1.377		1.606		1.946		4.277
		(1.300)		(2.089)		(1.310)		(4.179)
Ln(Dist-Pop-Weighted Vehicles p.c.)	0.741		0.605					
	(0.538)		(0.479)					
Ln(Dist-Pop-Weighted Vehicles p.c.)$_{t-1}$	0.150		0.111					
	(0.293)		(0.306)					
Ln(Dist-Pop-Weighted Vehicles p.c.)$_{t-2}$			-0.181					
			(0.432)					
Average of Ln(Dist-Pop-Weighted Vehicles p.c.) in t and $t-1$					0.864^{c}			
					(0.469)			
Average of Ln(Dist-Pop-Weighted Vehicles p.c.) in $t-1$ and $t-2$							0.494	
							(0.432)	
Ln(TFP)	0.864^{a}	-0.035	0.862^{a}	-0.213	0.916^{a}	-0.647	0.924^{a}	-2.807
	(0.042)	(1.136)	(0.052)	(1.817)	(0.032)	(1.221)	(0.036)	(3.900)
Ln(Av. Wages)	-0.344^{a}	-0.052	-0.350^{a}	0.009	-0.379^{a}	0.235	-0.391^{a}	1.166
	(0.043)	(0.453)	(0.050)	(0.746)	(0.033)	(0.494)	(0.037)	(1.619)
Observations	45,384	45,384	35,863	35,863	56,534	56,534	52,504	52,504
F-statistic	1.16		0.71		3.40		1.31	

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors, clustered at the municipality level in Panel A and at the province level in Panel B, are reported in parenthesis. All specifications include firm and sector-year fixed effects, as well as municipality-specific time trends. In Panel A, they additionally include province fixed effects.

database includes specific characters to allow for the identification of companies that are foreign entities or permanent establishments of entities not resident in Spain. The Bank of Spain also collects information on the Spanish firms that are linked with either foreign entities or permanent establishments of entities not resident in Spain and, thus, labels them as a multinational group. Data on single establishments and single manufacturing establishments is obtained from administrative tax data (AEAT, n.d.) reported by businesses in their tax returns of the Economic Activity Tax (*Impuesto de Actividades Económicas, IAE*). In particular, all businesses must report the geographic location of their plants as the taxable activity of each business is allocated among local jurisdictions according to the share of economic activity undertaken in each municipality

In columns 4 and 5 of Table G.5 we also explore the robustness of our intensive-margin results to alternative definitions of the "bust" period (2010-13 or 2011-13, instead of the baseline 2009-13).

As is clear from Table G.5, none of these robustness tests has a considerable effect on our estimates, with the second-stage elasticity ranging from -1.654 in column 1 to -2.118 in column 2.

Table G.5: Intensive-Margin Results with Alternative Samples

Sample:	Excl. Multinationals (1)	Single Manuf. Establishment (2)	Single Establishment (3)	Bust as 2010-2013 (4)	Bust as 2011-2013 (5)
OLS Elasticity	-0.269^a	-0.256^a	-0.427^a	-0.295^a	-0.318^a
	(0.028)	(0.041)	(0.067)	(0.033)	(0.033)
IV Elasticity	-1.654^a	-2.118^a	-1.751^a	-1.682^a	-1.703^a
	(0.247)	(0.339)	(0.387)	(0.286)	(0.301)
1st Stage Coeff.	1.358^a	1.225^a	1.744^a	1.241^a	1.231^a
	(0.120)	(0.139)	(0.335)	(0.124)	(0.124)
Observations	6,625	5,366	1,824	7,346	6,705
F-statistic	127.52	78.24	27.17	100.61	98.09

Notes: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province appear in parenthesis. All specifications include sector fixed effects.

G.4 Robustness to Adding Fixed Effects at Different Levels

In columns 1 and 2 of Table G.6, we replicate our baseline intensive margin results in Panel B of Table 3, and then proceed to demonstrate that the results are not materially affected by the inclusion of province fixed effects (columns 3 and 4) and province-sector fixed effects (columns 5 and 6). In Table G.7, we perform the same analysis (and reach the same conclusion) for the case of our specifications in Table 10 with total sales as the key right-hand side variable.

Table G.6: Intensive-Margin Results with Fixed Effects at Different Levels

Dependent Variable:	1st Stage (1)	2nd Stage (2)	1st Stage (3)	2nd Stage (4)	1st Stage (5)	2nd Stage (6)
ΔLn(Domestic Sales)		-1.607^a		-1.678^a		-1.572^a
		(0.248)		(0.237)		(0.232)
ΔLn(Dist-Pop-Weighted	1.312^a		1.394^a		1.457^a	
Vehicles p.c.)	(0.119)		(0.116)		(0.114)	
ΔLn(TFP)	1.023^a	2.810^a	1.027^a	2.874^a	1.030^a	2.774^a
	(0.028)	(0.213)	(0.029)	(0.205)	(0.030)	(0.203)
ΔLn(Av. Wages)	-0.526^a	-1.387^a	-0.528^a	-1.417^a	-0.526^a	-1.355^a
	(0.047)	(0.151)	(0.049)	(0.150)	(0.054)	(0.146)
Observations	8,009	8,009	8,009	8,009	7,821	7,821
Sector FE	Yes	Yes	Yes	Yes	No	No
Province FE	No	No	Yes	Yes	No	No
Sector-Province FE	No	No	No	No	Yes	Yes
F-statistic	122.44		144.25		164.65	

Notes: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Columns 1-2 replicate columns 4 and 8 of Table 3 (panel B) in the paper. Columns 3-6 re-estimate the same specification with different fixed effects.

Table G.7: Intensive-Margin Results with Total Sales adding Fixed Effects at Different Levels

Specification:	1st Stage (1)	2nd Stage (2)	1st Stage (3)	2nd Stage (4)	1st Stage (5)	2nd Stage (6)
ΔLn(Total Sales)		-2.374^a		-2.564^a		-2.380^a
		(0.526)		(0.529)		(0.511)
ΔLn(Dist-Pop-Weighted	0.888^a		0.913^a		0.962^a	
Vehicles p.c.)	(0.103)		(0.104)		(0.104)	
ΔLn(TFP)	1.063^a	3.690^a	1.064^a	3.879^a	1.066^a	3.691^a
	(0.026)	(0.482)	(0.027)	(0.487)	(0.031)	(0.470)
ΔLn(Average Wages)	-0.509^a	-1.750^a	-0.510^a	-1.840^a	-0.508^a	-1.737^a
	(0.043)	(0.250)	(0.044)	(0.257)	(0.047)	(0.242)
Observations	8,009	8,009	8,009	8,009	8,009	8,009
Sector FE	Yes	Yes	Yes	Yes	No	No
Province FE	No	No	Yes	Yes	No	No
Sector-Province FE	No	No	No	No	Yes	Yes
F-statistic	75.00		77.33		85.66	

Notes: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Columns 1-2 replicate columns 2 and 3 of Table 10 (panel B) in the paper. Columns 3-6 re-estimate the same specifications with different fixed effects.

G.5 Robustness to Clustering of Standard Errors at Different Levels

In this subsection, we analyze how our main results from Panel B of Tables 3 and 10 are affected by clustering the standard errors at different levels: province (baseline, as reported in the main text), municipality, two-way clustering by province and sector, and two-way clustering by municipality and sector.

The results in Table G.8 generalize those in panel B of Table 3, and the results in Table G.9 generalize those in panel B of Table 10. The standard errors are very similar to the baseline ones when we cluster at the municipality level, and somewhat larger when we use two-way clustering by province and sector and two-way clustering by municipality and sector. In all cases, the conclusions from our analysis are essentially unchanged. There are only a few coefficient estimates whose level of statistical significance shifts from under 1% to under 5% in Table G.8 and G.9.

Table G.8: Intensive-Margin 2SLS: Robustness to Different Levels of Clustering

Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔLn(Domestic Sales)					-10.068	-2.081	-1.751	-1.607
Level of Clustering:								
Province					(3.454)^a	(0.319)^a	(0.238)^a	(0.248)^a
Municipality					(3.462)^a	(0.254)^a	(0.204)^a	(0.208)^a
Prov. and sector					(4.482)^b	(0.610)^a	(0.414)^a	(0.410)^a
Munic. and sector					(4.511)^b	(0.602)^a	(0.410)^a	(0.400)^a
ΔLn(Dist-Pop-Wght. Vehicles p.c.)	0.339	1.194	1.346	1.312				
Level of Clustering:								
Province	(0.121)^a	(0.145)^a	(0.135)^a	(0.119)^a				
Municipality	(0.115)^a	(0.118)^a	(0.116)^a	(0.109)^a				
Prov. and sector	(0.147)^b	(0.210)^a	(0.170)^a	(0.139)^a				
Munic. and sector	(0.143)^b	(0.198)^a	(0.158)^a	(0.132)^a				
Observations	8,009	8,009	8,009	8,009	8,009	8,009	8,009	8,009
Control for TFP	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Control for Avg. Wages	No	No	Yes	Yes	No	No	Yes	Yes
Sector FE	No	No	No	Yes	No	No	No	Yes

Notes: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. This table presents estimates for the same regression specifications as in Table 3 (panel B) in the main text; thus, the point estimates coincide. We report the standard errors for the key covariate in the first-stage regressions (columns 1-4) and in the second-stage regressions (columns 5-8) assuming different levels of clustering of the standard errors: province (baseline), municipality, two-way cluster by province and sector, and two-way clustering by municipality and sector.

Table G.9: Intensive-Margin with Total Sales: Robustness to Different Levels of Clustering

Dependent Variable:	ΔLn(Exp) (1) OLS	ΔLn(TotSales) (2) 1st Stage	ΔLn(Exp) (3) 2nd Stage	ΔLn(TotSales) (4) 1st Stage	ΔLn(Exp) (5) 2nd Stage
ΔLn(Total Sales)	0.724		-2.374		-2.590
Level of clustering:					
Province	(0.050)[a]		(0.526)[a]		(0.606)[a]
Municipality	(0.038)[a]		(0.384)[a]		(0.416)[a]
Prov. and sector	(0.067)[a]		(0.909)[a]		(1.018)[b]
Munic. and sector	(0.060)[a]		(0.871)[a]		(0.961)[a]
ΔLn(Dist-Pop-Wght. Vehicles p.c.)		0.888		0.838	
Level of clustering:					
Province		(0.103)[a]		(0.107)[a]	
Municipality		(0.076)[a]		(0.076)[a]	
Prov. and sector		(0.194)[a]		(0.196)[a]	
Munic. and sector		(0.188)[a]		(0.188)[a]	
ΔLn(Stock of Capital)				0.101	0.382
Level of clustering:					
Province				(0.009)[a]	(0.067)[a]
Municipality				(0.008)[a]	(0.052)[a]
Prov. and sector				(0.011)[a]	(0.120)[a]
Munic. and sector				(0.011)[a]	(0.118)[a]
Observations	8,009	8,009	8,009	8,009	8,009
Control for TFP	Yes	Yes	Yes	Yes	Yes
Control for Avg. Wages	Yes	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes	Yes

Notes: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. This table presents estimates for the same regression specifications as in Table 10 (panel B) in the main text; thus, the point estimates coincide. We report the standard errors for the key covariate in the OLS regression (column 1), the first-stage regressions (columns 2 and 4) and the second-stage regressions (columns 3 and 5) assuming different levels of clustering of the standard errors: province (baseline), municipality, two-way cluster by province and sector, and two-way clustering by municipality and sector.

G.6 Weighted Least Squares Regressions

In this subsection, we analyze how our main results from panel B of Table 3 change if we use weighted least squares to estimate the regression parameters and weight firms according to different criteria: (i) by the log of average sales during the boom period (2002-2008); (ii) by the log of average employment during the boom period; (iii) by the log of the average assets during the boom period; and (iv) by the number of exporting years during the boom. The results in Table G.10 are very similar to those reported in panel B of Table 3, illustrating the robustness of those results.

G.7 Additional Controls

Table G.11 follows the same structure as Table 8 in the main text and considers some additional supply-side confounders, namely the firm-level average share of temporary workers during the boom

Table G.10: Intensive-Margin: Weighted Least Squares

Weighting Variable:	Ln(Avg. Sales) 2002-2008	Ln(Avg. Empl.) 2002-2008	Ln(Avg. Assets) 2002-2008	Nr. of Years Exporting
	(1)	(2)	(3)	(4)
OLS Elasticity	-0.289^a	-0.308^a	-0.289^a	-0.289^a
	(0.034)	(0.038)	(0.033)	(0.028)
IV Elasticity	-1.625^a	-1.629^a	-1.629^a	-1.796^a
	(0.250)	(0.230)	(0.249)	(0.287)
1st Stage Coefficient	1.281^a	1.240^a	1.274^a	1.190^a
	(0.120)	(0.129)	(0.123)	(0.142)
1st Stage F-Stat.	113.57	92.14	106.41	70.43
Observations	8,009	7,987	8,009	8,009

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province appear in parenthesis. All specifications include sector fixed effects.

(column 1), the boom-to-bust log change in the average number of bank offices in a firm's municipality (column 2), the boom-to-bust log change in the firm-level share of bank credit accounted for by short-term creditors (column 3), the interaction of that latter variable with the boom-to-bust log change in domestic sales (column 4), the boom-to-bust log change in the average number of bank offices in a firm's municipality (column 5), and the firm-level boom-to-bust log change in the number of permanent workers (column 6). The data on bank offices by municipality is from a proprietary dataset maintained by the Bank of Spain (Banco de España, n.d.*c*).

Table G.12 reports the results of estimating our main intensive-margin regression using the

Table G.11: Additional Controls

Dependent Variable:	ΔLn(Exports)					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Domestic Sales)	-1.561^a	-1.557^a	-1.642^a	-1.590^a	-1.651^a	-1.609^a
	(0.239)	(0.248)	(0.222)	(0.226)	(0.243)	(0.246)
Share of Temp. Workers in Boom	0.407^a					
	(0.109)					
ΔLn(Bank Offices p.c. in municipality)		-0.248^a				
		(0.079)				
ΔLn(Short-Term Creditors over Banking Credit)			-0.004	0.024		
			(0.011)	(0.041)		
ΔLn(Dom. Sales) ×Δ Ln(Short-Term Creditors over Banking Credit)				0.170		
				(0.215)		
ΔLn(Price of Land in munic.)					-0.030	
					(0.020)	
ΔLn(Permanent Workers)						0.001
						(0.010)
Observations	7,889	7,933	6,429	6,429	7,300	7,631
F-statistic	117.26	134.56	89.61	89.14	104.12	127.10

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province appear in parenthesis. All specifications include sector fixed effects. Same specifications as Table 8.

gravity-based instrument including controls for firm size in the Boom period (2002-2008), where size could be measured either by the average annual sales (column 1), average annual employment (column 2) or average annual assets (column 3). We also control for the firms' exports-to-sales ratio in the boom period (column 4). The IV results are generally similar to those obtained in our baseline specification (see panel B of Table 3 in the paper).

Table G.12: Intensive-Margin 2SLS with Additional Controls

Additional Control:	Ln(Avg. Sales) 2002-2008	Ln(Avg. Empl.) 2002-2008	Ln(Avg. Assets) 2002-2008	Exports-to-Sales 2002-2008
	(1)	(2)	(3)	(4)
OLS Elasticity	-0.286^a	-0.288^a	-0.284^a	-0.199^a
	(0.033)	(0.033)	(0.032)	(0.034)
IV Elasticity	-1.432^a	-1.531^a	-1.473^a	-1.829^a
	(0.203)	(0.225)	(0.209)	(0.360)
1st Stage Coefficient	1.595^a	1.476^a	1.548^a	0.992^a
	(0.121)	(0.122)	(0.121)	(0.121)
1st Stage F-Stat.	173.06	146.86	163.76	67.74
Observations	8,009	8,009	8,009	8,009

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province appear in parenthesis. All specifications include sector fixed effects.

G.8 Further Robustness Tests Related to Gravity-Based Instrument

In this section, we report additional robustness tests related to the gravity-based instrument. We test how the results vary when we restrict the sampple to firms that already exported to multiple markets in the boom period, estimate the gravity equations at the province level, and report our main intensitve-margin results using different gravity models to construct the weighted instrument.

Table G.13 presents estimates analogous to those in column 4 of Table 1 but computed using information exclusively on subsamples of municipalities where firms that exported to a minimum number of foreign markets at least once during the boom period are located. For the sake of facilitating the comparison, we present in column 1 of Table G.13 the same estimates presented in column 4 of Table 1. These are computed using information on all firms that exported at least once during the boom and bust periods and that were active in 2006, the only year for which the the Spanish Tax Agency (*Agencia Estatal de Administración Tributaria*, AEAT) has provided us with information on the municipality of destination of firms' domestic sales.

As a way to evaluate the possible bias coming from the large presence of zeroes in our firm-to-municipality matrix of flows, we present in columns 2 to 5 of Table G.13 results for subsamples of firms that are likely to be less affected by this source of bias. Specifically, to compute the estimates in columns 2, 3, 4, and 5 of Table G.13, we restrict the sample to firms that exported to at least 2, 3, 4 and 5 distinct countries in the boom period, respectively. The elasticities with respect to distance and population remain roughly constant as we restrict the sample to firms that exported to a larger number of destinations. In fact, in results available upon request, we observe that, when recomputing the results in Panel B of Table 3 for the subsamples of firms used to compute the estimates in columns 2 to 5 of Table G.13, we consistently obtain TSLS estimates of the elasticity of export flows with respect to domestic sales that are close to the baseline estimates reported in Table 3.

Table G.13: Estimates from Gravity Equations For Subsamples of Firms by Number of Export Destinations in Boom

Dependent Variable:	Ln(Firm-to-Municipality Trade Flows)				
	(1)	(2)	(3)	(4)	(5)
Number Export Destinations:	≥ 1	≥ 2	≥ 3	≥ 4	≥ 5
Ln(Distance)	-0.150^a	-0.148^a	-0.147^a	-0.142^a	-0.143^a
	(0.021)	(0.020)	(0.021)	(0.021)	(0.021)
Ln(Population)	0.300^a	0.306^a	0.312^a	0.320^a	0.327^a
	(0.012)	(0.013)	(0.014)	(0.014)	(0.015)
Observations	675,715	607,755	550,281	487,818	443,134
R-squared	0.18	0.18	0.19	0.19	0.19
Municipality-Origin FE	Yes	Yes	Yes	Yes	Yes

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered at the province (of origin) level are reported in parenthesis. The data on province-level trade flows for manufacturing firms is for the 2006 fiscal year. Ln(Population) denotes the log of the population of the destination province in 2006. Ln(Distance) denotes the log of the distance, in kilometers, between the two provinces in each pair. The estimates in column 1 correspond to the baseline estimates in column 4 of Table 1. The estimates in columns 2, 3, 4, and 5 restrict the sample to firms that export to at least 2, 3, 4 and 5 distinct countries in the boom period, respectively.

Table G.14 reports results from estimating a gravity equation aggregating our data at the province level, instead of at municipality level as in Table 1 in the paper. Specifically, columns 1 and 2 report results using province-to-province sales data, while columns 3 and 4 reports results using firm-to-province sales data. The point estimate for the coefficient on log distance is close to -1 in all four specifications and the coefficient on log population is approximately 1.3 in all specifications, both highly significant. Columns 2 and 4 also include an own-province dummy to capture home bias, with estimated coefficients equal to 1.45 and 1.34, respectively. These results suggest that the degree of home bias in our data for Spain is consistent with the literature, even though the point estimates for the coefficients on log distance and log population are smaller in absolute value in our municipality-level regressions reported in Table 1.

Table G.15 reports estimates of our main intensive-margin specifications, but constructing the gravity-based instrument using the estimates from column 2 of Table 1 (which includes own-municipality and own-province dummies) instead of those of column 1 of Table 1 (which do not include those dummies). Specifically, we include the estimated coefficient of the own-province dummy when constructing the weights for the instrument, but when computing the weight we do not include the own-municipality dummy to avoid the potential endogeneity of local sales. Columns 1-4 of Table G.15 report first-stage estimates and column 5-8 report second-stage estimates, following the same structure as Panel B of Table 3. The results are broadly consistent with those reported in the paper, with a second-stage elasticity of exports with respect to changes in (instrumented) domestic sales equal to -1.614 in column 8 (compared to -1.607 in the same column of Table 3, Panel B). Therefore, we conclude that our results are robust to using alternative weights to construct the gravity-based instrument that account for home bias at the province level.

Table G.16 is a variant of Table G.15 where, instead of using the point estimates from our gravity equation, we impose "standard" coefficients on log distance, log population and the own-province dummy. Specifically, we construct the weights for the gravity-based instrument using $\beta_{pop} = 1$, $\beta_{dist} = -1$, and $\beta_{ownprov} = 1$. The table follows the same structure as Table G.15. The first-stage coefficients in columns 1-4 are smaller in absolute value but still significant, with F-statistics between 12.8 and 15.9 for the specifications with controls. The second-stage coefficients

Table G.14: Estimates from Gravity Equations at Provincial Level

Dependent Variable:	Ln(Bilateral Trade Flows between Provinces)			
	(1)	(2)	(3)	(4)
Ln(Distance)	-1.225[a]	-1.091[a]	-1.003[a]	-0.877[a]
	(0.050)	(0.049)	(0.047)	(0.048)
Ln(Population)	1.346[a]	1.332[a]	1.391[a]	1.378[a]
	(0.028)	(0.028)	(0.023)	(0.023)
Dummy for own-province flows		1.449[a]		1.344[a]
		(0.150)		(0.130)
Observations	2,597	2,597	2,545	2,545
R-squared	0.83	0.84	0.85	0.85
Province-Origin FE	Yes	Yes	Yes	Yes

Note: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered at the province (of origin) level are reported in parenthesis. The data on province-level trade flows for manufacturing firms is for the 2006 fiscal year. Ln(Population) denotes the log of the population of the destination province in 2006. Ln(Distance) denotes the log of the distance, in kilometers, between the two provinces in each pair. The estimates in columns 1 to 2 use province-to-province sales data; the estimates in columns 3 and 4 use firm-to-province data.

Table G.15: Intensive-Margin Results using Gravity IV from column 2 in Table 1

Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ Ln(Domestic Sales)					-10.143[a]	-2.082[a]	-1.752[a]	-1.608[a]
					(3.453)	(0.316)	(0.235)	(0.245)
Δ Ln(Dist-Pop-Weighted Vehicles p.c.)	0.339[a]	1.201[a]	1.353[a]	1.322[a]				
	(0.120)	(0.144)	(0.133)	(0.117)				
Δ Ln(TFP)		0.829[a]	1.031[a]	1.024[a]		2.624[a]	2.877[a]	2.811[a]
		(0.028)	(0.029)	(0.028)		(0.239)	(0.220)	(0.211)
Δ Ln(Average Wages)			-0.621[a]	-0.526[a]			-1.621[a]	-1.387[a]
			(0.037)	(0.047)			(0.173)	(0.151)
Observations	8,009	8,009	8,009	8,009	8,009	8,009	8,009	8,009
Sector FE	No	No	No	Yes	No	No	No	Yes
F-statistic	8.01	69.76	103.11	128.04				

Note: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered by province appear in parenthesis. For any X, ΔLn(X) is the log difference between the average of X in 2009-2013 and its average in 2002-2008. ΔLn(Dist-Pop-Weighted vehicles p.c.) is the instrument constructed using data on vehicles per capita at the municipal level and applying the weights from the gravity equation reported in column 2 of Table 1. Columns 1-4 contain first-stage estimates; columns 5-8 contain second-stage estimates. F-statistic denotes the corresponding test statistic for the null hypothesis that the coefficient on Ln(Dist-Pop-Weighted Vehicles p.c.) equals zero.

on the (instrumented) change in log domestic sales in columns 5-8 are somewhat smaller in absolute value, generally close to -1.

Table G.17 presents another variant of Table G.16, where we impose the following coefficients to construct the weights of the gravity-based instrument: $\beta_{pop} = 1$, $\beta_{dist} = -1$, and $\beta_{ownprov} = 1.5$. The results are very similar to those of Table G.16, although in this case the point estimate in column 8 is -0.866 and it is no longer statistically significant at conventional levels.

Table G.16: Intensive-Margin Results with Fixed Coefficients in the Gravity Equation (I)

Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ Ln(Domestic Sales)					-2.974^a	-1.212^b	-1.207^a	-0.964
					(1.116)	(0.524)	(0.461)	(0.582)
Δ Ln(Dist-Pop-Weighted Vehicles p.c.)	0.137^a	0.236^a	0.237^a	0.208^a				
	(0.047)	(0.060)	(0.062)	(0.058)				
Δ Ln(TFP)		0.796^a	0.987^a	0.974^a		1.936^a	2.341^a	2.187^a
		(0.026)	(0.028)	(0.028)		(0.399)	(0.422)	(0.522)
Δ Ln(Average Wages)			-0.603^a	-0.505^a			-1.292^a	-1.062^a
			(0.038)	(0.050)			(0.260)	(0.266)
Observations	8,009	8,009	8,009	8,009	8,009	8,009	8,009	8,009
Sector FE	No	No	No	Yes	No	No	No	Yes
F-statistic	8.61	15.79	14.51	12.88				

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province appear in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the log difference between the average of X in 2009-2013 and its average in 2002-2008. The weights for the gravity-based instrument are constructed using the fixed coefficients $\beta_{pop} = 1$, $\beta_{dist} = -1$, and $\beta_{ownprov} = 1$.

Table G.17: Intensive-Margin Results with Fixed Coefficients in the Gravity Equation (II)

Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ Ln(Domestic Sales)					-2.508^b	-1.080^c	-1.098^b	-0.846
					(1.082)	(0.579)	(0.524)	(0.669)
Δ Ln(Dist-Pop-Weighted Vehicles p.c.)	0.114^a	0.185^a	0.183^a	0.159^a				
	(0.038)	(0.046)	(0.048)	(0.045)				
Δ Ln(TFP)		0.794^a	0.986^a	0.973^a		1.831^a	2.235^a	2.072^a
		(0.026)	(0.028)	(0.028)		(0.443)	(0.485)	(0.610)
Δ Ln(Average Wages)			-0.603^a	-0.505^a			-1.226^a	-1.002^a
			(0.039)	(0.050)			(0.297)	(0.312)
Observations	8,009	8,009	8,009	8,009	8,009	8,009	8,009	8,009
Sector FE	No	No	No	Yes	No	No	No	Yes
F-statistic	9.05	16.10	14.67	12.66				

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province appear in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the log difference between the average of X in 2009-2013 and its average in 2002-2008. The weights for the gravity-based instrument are constructed using the fixed coefficients $\beta_{pop} = 1$, $\beta_{dist} = -1$, and $\beta_{ownprov} = 1.5$.

G.9 Regressions at the Municipality-Sector Level

We report here results of regressions where all variables other than the instrument (defined at the municipality level) are defined as the average across firms within a municipality and a sector. Aggregating at the municipality-sector level, and not simply at the municipality level, as, consistently with our baseline specification, it allows us to include sector fixed effects.

Table G.18 reports OLS estimates: the coefficient on domestic sales is positive and significant when not controlling for supply factors (TFP, wages), and negative and significant when doing so.

Table G.18: Regressions at Municipality-Sector Level: Ordinary Least Squares

Dependent Variable:	ΔLn(Exports)			
	(1)	(2)	(3)	(4)
ΔLn(Domestic Sales)	0.310[a]	0.317[a]	0.332[a]	-0.267[a]
	(0.045)	(0.044)	(0.042)	(0.026)
ΔLn(TFP)				1.642[a]
				(0.055)
ΔLn(Average Wages)				-0.795[a]
				(0.070)
Observations	4,080	4,080	4,080	4,080
R-squared	0.037	0.058	0.094	0.264
Province FE	No	Yes	Yes	Yes
Sector FE	No	No	Yes	Yes

Notes: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered by province appear in parenthesis.

Table G.19 reports the IV results for the same specifications considered in Table G.18. The instrument is weak when no supply controls are included. Controlling for supply factors makes the instrument strong. The TSLS coefficient in this case is -1.02, smaller in magnitude than that in column 8 of Table 3 (panel B), but broadly comparable and statistically significant.

Table G.19: Regressions at Municipality-Sector Level: Two-Stage Least Squares

Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔLn(Domestic Sales)					-13.261	-13.064	-11.612[c]	-1.532[a]
					(9.805)	(8.993)	(6.686)	(0.370)
ΔLn(Dist-Pop-Weighted Vehicles p.c.)	0.230	0.263	0.292	1.423[a]				
	(0.183)	(0.189)	(0.177)	(0.136)				
ΔLn(TFP)				1.198[a]				3.123[a]
				(0.029)				(0.411)
ΔLn(Avg. Wages)				-0.500[a]				-1.410[a]
				(0.059)				(0.233)
Observations	4,080	4,080	4,080	4,080	4,080	4,080	4,080	4,080
Province FE	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Sector FE	No	No	Yes	Yes	No	No	Yes	Yes
F-statistic	1.58	1.93	2.72	109.47				

Notes: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered by province appear in parenthesis.

H Alternative Identification Strategies

H.1 Results Exploiting Export Demand Shocks in 2002-07

We implement here an exercise analogous to that in Berman, Berthou and Héricourt (2015) with the aim of determining the sign of the impact of demand-driven changes in exports on firms' domestic sales. More specifically, denoting as R_{ixt} and R_{idt} the total foreign and domestic sales of firm i in year t, respectively, and denoting as R_{it} the total sales if i in t (i.e., $R_{it} \equiv R_{ixt} + R_{idt}$), we present here OLS and TSLS (two-stage least squares) estimators of the elasticity of R_{idt} with respect to changes in either R_{ixt} or R_{it} that are driven by export demand shocks. To compute these TSLS estimates, we use a shift-share instrument analogous to that in Berman, Berthou and Héricourt (2015); more specifically, our instrumental variable is:

$$Z_{it} \equiv \sum_j \omega_{ij} M_{jt}, \tag{H.1}$$

where ω_{ij} denotes the share of each destination j in firm i's exports over the sample period 2002-2007, and M_{jt} denotes the total imports (excluding imports from Spain) of destination j in year t. As controls, we include year and firm fixed effects and, in some specifications, also firm-specific time trends.

Additionally, we also present OLS and TSLS estimators of the elasticity of the one-year change in R_{idt} with respect to the one-year change in either R_{ixt} or R_{it}. To compute these TSLS estimates, we use the shift-share instrument

$$\Delta Z_{it} \equiv \sum_j \omega_{ij} \Delta M_{jt}, \tag{H.2}$$

where ω_{ij} is defined as in equation (H.1) and, for every random variable X_{it}, $\Delta X_{it} \equiv X_{it} - X_{it-1}$. In these first-difference specifications, we include year fixed effects and, in some of them, also firm fixed effects.

We present our results in Table H.1. Panel A presents OLS estimates, while Panel B presents TSLS estimates. The specifications whose estimates are reported in columns 1 to 4 use the log of exports as the key covariate of interest; those reported in columns 5 to 8 use the log of total sales instead. Columns 1, 2, 5, and 6 present estimates for specifications in levels; columns 3, 4, 7, and 8 do so for specifications in first differences. While all specifications other than those columns 3 and 7 include firm fixed effects, only those in columns 2 and 6 incorporate a firm-specific time trend in the regression.

The information on export flows by firm and destination country is only available for the period 2002-2007.[6] Thus, we compute the weight ω_{ij} for every firm i and destination j using the 2002-2007 export data, and estimate the elasticity of domestic sales with respect to either aggregate exports or total sales using only information for this sample period. To construct the instruments Z_{it} and ΔZ_{it} defined in equations (H.1) and (H.2), we combine our measures of ω_{ij} for every destination j and every firm i in the sample with measures of M_{jt} and ΔM_{jt} constructed using the information on country and year-specific imports reported in the UN Comtrade dataset (United Nations, n.d.).

Contrary to the findings in Berman, Berthou and Héricourt (2015), and consistently with the estimates we present in section IV in the main draft, our OLS estimates of the elasticity domestic sales with respect to aggregate exports are negative (see Panel A, columns 1 to 4, in Table H.1). When using the shift-share instruments described in equations (H.1) and (H.2) to compute TSLS estimates of the elasticity domestic sales with respect to aggregate exports (see Panel B, columns

[6] After 2007, a change in the methodology renders the information on export flows by destination country unreliable.

Table H.1: Regressions à la Berman, Berthou and Héricourt (2015)

Specification:	Level		First-difference		Level		First-difference	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Panel A: OLS estimator				
Exports	-0.051^a	-0.135^a	-0.125^a	-0.149^a				
	(0.009)	(0.010)	(0.010)	(0.012)				
Total sales					1.002^a	1.140^a	1.212^a	1.284^a
					(0.027)	(0.047)	(0.046)	(0.057)
Firm fixed effect	Yes	Yes	No	Yes	Yes	Yes	No	Yes
Firm spec. time trend?	No	Yes	No	No	No	Yes	No	No
Obs.	24,142	24,142	19,023	19,023	24,142	24,142	19,023	19,023
				Panel B: TSLS estimator				
Exports	0.046	-0.373^a	-0.050^c	-0.093^a				
	(0.075)	(0.072)	(0.027)	(0.035)				
Total sales					0.313	-6.810^b	-0.559	-1.310^b
					(0.503)	(2.698)	(0.340)	(0.627)
Firm fixed effect	Yes	Yes	No	Yes	Yes	Yes	No	Yes
Firm spec. time trend?	No	Yes	No	No	No	Yes	No	No
1st. Stage F-test	332.70	315.30	456.28	350.20	44.10	9.97	60.97	30.92
Obs.	24,142	24,142	19,023	19,023	24,142	24,142	19,023	19,023

Notes: ^{a} denotes 1% significance; ^{b} denotes 5% significance; ^{c} denotes 10% significance. All specifications include year fixed effects.

1 to 4, in Table H.1), we obtain estimates that are either non-statistically different from zero (in column 1) or negative (in columns 2 to 4).

As discussed in section VI.A in the main draft and in Appendix E.2, a model with increasing marginal costs predicts a constant elasticity of domestic sales with respect to total sales. We estimate such elasticity in columns 5 to 8 in Table H.1. In this case, the positive sign of the OLS estimates reported in Panel A is not very informative about the slope of the firm's marginal cost curve, as supply shocks (e.g., productivity, factor prices) make a firm's total sales and domestic sales positively correlated. Conversely, the TSLS estimates reported in Panel B have the potential of being informative about the causal effect of export-demand-driven changes in total sales on a firm's domestic sales. These TSLS estimates are not statistically different from zero in columns 1 and 3; however, they become negative and statistically different from zero when firm-specific time trends are allowed in the regression specification in levels (see column 2) or, equivalently, when firm fixed effects are allowed in the regression specification in first differences.

In sum, when running regressions à la Berman, Berthou and Héricourt (2015), we find no evidence supporting the positive causal relationship between exports and domestic sales that these authors previously found. We find that this relationship is either negative or not statistically different from zero. The data used in Berman, Berthou and Héricourt (2015) contains information on firms located in a different country (France) and in a different sample period (1995-2001). Differences either in the predictability of the export demand shocks used as instruments or in the share of firms operating in capacity-constrained sectors could explain the differences between the estimates presented in Table H.1 and those presented in Berman, Berthou and Héricourt (2015); a study of the precise causes behind these differences is beyond the scope of this paper.

H.2 Alternative Instruments: the Primitive Causes of the Spanish Crisis

The instrumental variable estimates presented in this paper use as instruments proxies for the change in demand faced by firms in a municipality and, thus, do not require taking a stand on the source or cause of the observed demand changes. We next construct alternative instruments that attempt to better capture the deep roots of the Great Recession in Spain. For each of these instruments, whenever a measure at the municipality level is available, we always weigh these municipality-level demand shocks faced by firms in a given municipality in the same manner as we did for our baseline gravity-based instrument.

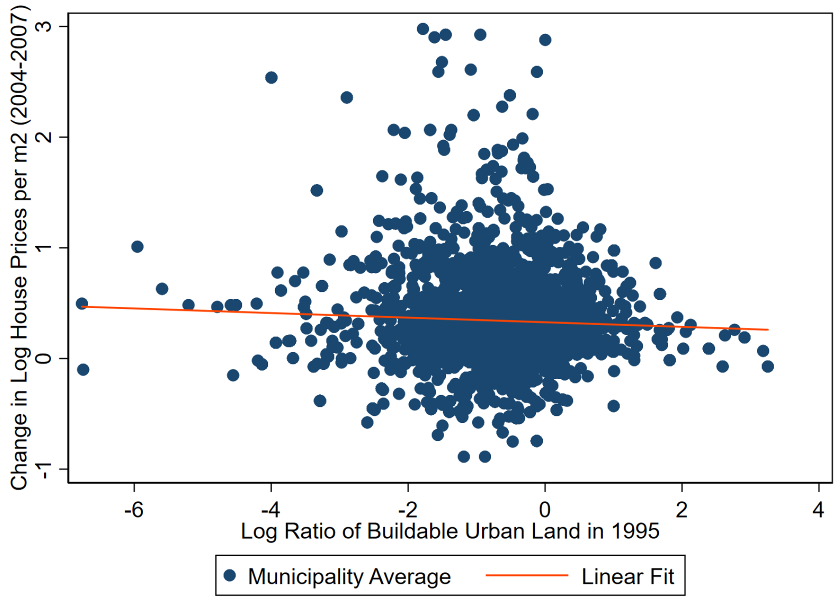
As described in section II.A of the main text, the Great Recession in Spain was largely driven by a real estate bubble. Our first alternative instrument thus attempts to identify an exogenous source of the intensity of the bubble across different locations. More precisely, we construct ratios of available 'buildable' urban land to urban land with already built structures in the year 1995, a year sufficiently removed from the housing boom (for detail on these data, see Basco, Lopez-Rodriguez and Elias, 2020). We conjecture that this ratio is a proxy for the housing supply elasticity in a given municipality, and that municipalities with lower housing supply elasticities should have experienced larger housing price increases during the boom years and, as a result, larger reductions in household wealth and consumption during the bust years. Indeed, as we show in Figure H.1, there is a negative cross-sectional correlation between these housing supply elasticities (proxied by the 1995 ratio of available 'buildable' urban land to urban land with already built structures) and housing price growth during the boom years 2004-07. This alternative instrumentation strategy is however not without limitations: a potential threat to its validity is the fact that housing supply elasticities could also operate as shifters of the firm's marginal costs, by affecting the cost of non-residential structures (i.e., factories).[7]

Our second alternative instrument is motivated by the importance of tourism revenue for the Spanish economy. Driven by the drop in demand in foreign countries, the number of foreign tourists visiting Spain peaked in 2007 at 58.66 millions visitors, before falling by more than 10% to 52.18 million and 52.68 million visitors in 2009 and 2010, respectively. Because tourism revenue accounts for roughly 10% of Spanish GDP, and because the decline in foreign visitors affected different regions in Spain differently, this generates an alternative source of geographical variation in local demand. We use a 2001 province-specific measure of exposure to tourism shocks, interacted with the log change in tourists at the national level between the boom and the bust, as an instrument for the boom-to-bust changes in demand in the corresponding province. Our measure of exposure is the number of foreign tourists that visited a province in 2001 divided by the population of the province in the same year.

We finally develop a third set of alternative instruments related to the construction sector. The burst of the real estate bubble affected directly the construction sector. As mentioned in footnote 10 in the main text, the share of total employment in the construction sector peaked at 13.5% in the summer of 2007 and then collapsed, reaching 5.4% by early 2014. A large share of the workers employed in the construction sector during the boom ended up unemployed during the bust period. These workers saw their consumption capacity severely reduced in the bust period relative to the boom. Consequently, one may conjecture that the boom-to-bust drop in demand for manufacturing products was larger in those municipalities for which the construction sector was a particularly important source of income during the boom years. Accordingly, we use the 2001 construction wage bill share in a municipality, interacted with the log change in the national construction wage bill between the boom and the bust, as a determinant of the boom-to-bust

[7] More specifically, municipalities with a lower housing supply elasticity might have experienced larger boom-to-bust reductions in the cost of land, which might have contributed to a larger relative export growth for firms located in those municipalities.

Figure H.1: Housing Supply Elasticities and Housing Price Growth during 2004-07



changes in demand in the corresponding municipality.[8] We further explore the robustness of this instrument to alternatives using municipality-level log changes in employment and turnover in the construction sector rather than log changes in the sector's wage bill.

In Table H.2, we report the results obtained under these different alternative instruments. Although the first-stage F-test statistics associated with two of these instruments are below ten and, thus, one should be cautious interpreting the corresponding second-stage estimates, it is worth remarking that the second-stage elasticities of exports to domestic sales are all quite similar to those obtained with our benchmark instrumentation strategy in Table 3. Furthermore, the p-values of the Sargan test of overidentifying restrictions are generally quite large and do not reject the validity of our baseline instrument. In sum, these results enhance our confidence in the existence of a causal relationship between demand-driven changes in domestic sales shocks and changes exports, with an elasticity roughly equal to -1.6.

In terms of data sources, the data to construct the proxy for the housing supply elasticity in a given municipality come from the Spanish Cadastre (Dirección General del Catastro). In particular, we use the measure developed by Basco, Lopez-Rodriguez and Elias (2020) which is a municipality-specific ratio of available "buildable" urban land to urban land with already built structures. The municipality-specific residential house prices used in Figure H.1 are obtained from the census of real-estate transactions owned by the Spanish Ownership Registry (Registro de la Propiedad, n.d.). We calculate the market value price per square meter for each residential housing transaction and

[8] The relevance and validity of our instrument does not depend on the fact that we multiply the municipality-specific 2001 construction wage bill share by the boom-to-bust log change in the national construction wage bill, which is common to all observations in our regression. We introduce this shifter in our instrument for the sake of facilitating the interpretation of the first-stage coefficient on this instrument. When interpreting our results, one should bear in mind that identification must come then from assumptions imposed on the distribution of the 2001 construction wage bill. See Goldsmith-Pinkham, Sorkin and Swift (2020) for a discussion of identification in this context.

Table H.2: Additional Alternative Instruments and Overidentification Tests

Dependent Variable:	ΔLn(Domestic Sales)				
	(1)	(2)	(3)	(4)	(5)
Ln(Urban Land Supply Ratio in 1995)	0.197^{b}				
(Weighted by Distance and Population)	(0.098)				
ΔLn(foreign tourists) ×		0.256^{a}			
2001 foreign tourists p.c. in prov.		(0.092)			
ΔLn(construction wage bill) ×			0.381^{a}		
2001 wage bill share in munic.			(0.062)		
ΔLn(construction employment) ×				0.428^{a}	
2001 empl. share in munic.				(0.074)	
ΔLn(construction turnover) ×					0.160^{a}
2001 turnover share in munic.					(0.025)

Dependent Variable:	ΔLn(Exports)				
ΔLn(Domestic Sales)	-1.401^{b}	-1.023^{a}	-1.229^{b}	-0.875	-1.533^{a}
	(0.634)	(0.286)	(0.532)	(0.626)	(0.524)
Observations	8,009	8,009	7,935	7,935	7,935
F-statistic	4.04	7.66	37.28	33.49	40.59
P-value for Sargan test	0.80	0.10	0.78	0.25	0.93

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors are clustered by province in columns 1-2 and by clustered by municipality in columns 3-5. All specifications include firm-level log TFP and log wages as additional controls. All specifications include sector fixed effects.

then aggregate those prices for all transactions made in a municipality during a natural year to create yearly average prices per square meter. The data on the number of foreign tourists at the province level come from the Spanish National Statistical Office (INE, n.d.*b*). Finally, the wage bill, employment and turnover in the construction sector are computed based on our data from the Commercial Registry (*Registro Mercantil Central*).

I Regression Results with Total Sales instead of Domestic Sales

We present here regression estimates for specifications analogous to those in Tables 3 to 9 in the main draft for the gravity-based instrument, with the only difference that the boom-to-bust log change in *total* sales is included as a right-hand-side variable instead of the log change in *domestic* sales.

Table I.1 replicates panel B of Table 3, with total sales instead of domestic sales. Note that columns 4 and 8 of are identical to columns 2 and 3 in panel B of Table 10.

Table I.1: Intensive Margin: Two-Stage Least Squares Estimates

Dependent Variable:	ΔLn(Total Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔLn(Total Sales)					20.278^c	-3.446^a	-2.724^a	-2.374^a
					(11.429)	(0.848)	(0.571)	(0.526)
ΔLn(Dist-Pop-Weighted Vehicles p.c.)	-0.168	0.721^a	0.865^a	0.888^a				
	(0.100)	(0.123)	(0.108)	(0.103)				
ΔLn(TFP)		0.862^a	1.054^a	1.063^a		3.869^a	3.942^a	3.690^a
		(0.021)	(0.023)	(0.026)		(0.681)	(0.541)	(0.482)
ΔLn(Average Wages)			-0.590^a	-0.509^a			-2.139^a	-1.750^a
			(0.037)	(0.043)			(0.332)	(0.250)
F-statistic	2.81	34.37	64.32	75.00				
Observations	8,009	8,009	8,009	8,009	8,009	8,009	8,009	8,009
Sector FE	No	No	No	Yes	No	No	No	Yes

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province appear in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the log difference between the average of X in 2009-2013 and its average in 2002-2008. Columns 1 to 4 contain first-stage estimates; columns 5 to 8 contain second-stage estimates. F-statistic denotes the corresponding test statistic for the null hypothesis that the coefficient on Ln(Dist-Pop-Weighted Vehicles p.c.) equals zero.

Table I.2 replicates Table 4, with total sales on the right-hand side. The second-stage coefficient in specifications with 3- and 2-year rolling averages (columns 3 and 6, respectively) are comparable to those in Table 4, although larger in absolute value. In specifications with yearly data (columns 7-9), both instruments are weak and the 2SLS results are insignificant.

Table I.3 replicates Table 5, with total sales on the right-hand side. We find that the sign of the effect of domestic demand shocks on the extensive margin of exports is sensitive to the specific way in which this extensive margin is measured, just as we found for regressions in which the key right-hand-side variable was domestic sales.

Table I.2: Panel Regressions

	Panel A: Municipality-level Instrument								
Data Frequency:	3-year Moving Average			2-year Moving Average			Annual Data		
	OLS (1)	1st Stage (2)	2SLS (3)	OLS (4)	1st Stage (5)	2SLS (6)	OLS (7)	1st Stage (8)	2SLS (9)
Ln(Total Sales)	0.832[a]		-2.915[b]	0.869[a]		-2.437	0.919[a]		-5.000
	(0.030)		(1.371)	(0.028)		(1.857)	(0.030)		(20.487)
Ln(Vehicles p.c. in municipality)		0.138[a]			0.082[a]			-0.010	
		(0.040)			(0.032)			(0.035)	
Ln(TFP)	0.358[a]	0.989[a]	4.059[a]	0.311[a]	0.987[a]	3.574[c]	0.237[a]	0.975[a]	6.010
	(0.048)	(0.021)	(1.354)	(0.045)	(0.021)	(1.839)	(0.046)	(0.023)	(19.978)
Ln(Average Wages)	-0.154[a]	-0.442[a]	-1.808[a]	-0.140[a]	-0.429[a]	-1.556[c]	-0.106[a]	-0.405[a]	-2.506
	(0.042)	(0.023)	(0.617)	(0.038)	(0.022)	(0.807)	(0.033)	(0.022)	(8.295)
Observations	66,711	66,710	66,710	65,709	65,708	65,708	60,199	60,198	60,198
F-statistic		12.01			6.70			0.08	
	Panel B: Distance- and population-weighted Instrument								
Data Frequency:	3-year Moving Average			2-year Moving Average			Annual Data		
	OLS (1)	1st Stage (2)	2SLS (3)	OLS (4)	1st Stage (5)	2SLS (6)	OLS (7)	1st Stage (8)	2SLS (9)
Ln(Total Sales)	0.832[a]		-3.589[a]	0.869[a]		-4.429[a]	0.916[a]		2.041[c]
	(0.039)		(0.610)	(0.036)		(1.306)	(0.040)		(1.173)
Ln(Dist-Pop-Weighted Vehicles p.c.)		0.611[a]			0.417[a]			0.685[c]	
		(0.076)			(0.096)			(0.357)	
Ln(TFP)	0.358[a]	0.995[a]	4.725[a]	0.311[a]	0.988[a]	5.540[a]	0.236[a]	0.976[a]	-0.862
	(0.037)	(0.019)	(0.566)	(0.040)	(0.021)	(1.232)	(0.051)	(0.024)	(1.158)
Ln(Average Wages)	-0.154[a]	-0.446[a]	-2.105[a]	-0.140[a]	-0.429[a]	-2.409[a]	-0.105[a]	-0.406[a]	0.352
	(0.033)	(0.027)	(0.259)	(0.031)	(0.026)	(0.527)	(0.029)	(0.028)	(0.469)
Observations	66,711	66,711	66,711	65,709	65,709	65,709	60,199	60,199	60,199
F-statistic		65.40			18.75			3.68	

Note: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors, clustered at the municipality level in Panel A and at the province level in Panel B, are reported in parenthesis. All specifications include firm and sector-year fixed effects, as well as municipality-specific time trends. In Panel A, they additionally include province fixed effects. The dataset used in columns 1-3 is constructed calculating three-year moving averages of all the variables for each firm, where the periods are 2002-2004, 2003-2005, etc., for a total of ten periods. In columns 4-6, we calculate two-year moving averages, where the periods are 2002-2003, 2003-2004, etc., for a total of 11 periods. If a firm is missing from the data for one year, the moving average including that year is calculated only based on the existing observations. In columns 7-9, we use the original annual data with 12 periods between 2002 and 2013.

Tables I.4 to I.7 replicate the robustness tests presented in Tables 6 to 9 in the main draft when using the gravity-based instrument, again with the only difference of including total sales instead of domestic sales on the right-hand side. Concerning Table I.4, excluding firms related to the auto industry by geographic proximity (panels A-C) does not have a significant impact on the main elasticity coefficient in the second-stage regression, always around -2.5 (columns 3 and 6 in the top panel and column 3 in the bottom panel). Excluding industries that are among the top-two suppliers or clients of the auto industry (panel D) does make the elasticity coefficient larger in absolute value, reaching -3.188.

Table I.5 reports the results from specifications where we use alternative instruments, as in Table 7 in the paper. In columns 1 to 4, we test the robustness of the results in Table I.1 to instruments constructed under alternative specifications of the gravity equation. Column 1 reproduces the baseline estimates (see columns 4 and 8 in Table I.1). In column 2, we include own-municipality and own-province dummies in the gravity equation we use to compute the estimated distance

Table I.3: Extensive Margin: Two-Least Squares Estimates

Dependent Variable:		Export Dummy		Proportion of Years	
	1st Stage	OLS	2nd Stage	OLS	2nd Stage
	(1)	(2)	(3)	(4)	(5)
Ln(Total Sales)		0.067^a	-0.159^a	0.052^a	0.049^b
		(0.007)	(0.041)	(0.006)	(0.020)
Ln(Dist-Pop-Weighted Vehicles p.c.)	0.929^a				
	(0.113)				
Ln(TFP)	1.195^a	0.016^b	0.279^a	0.014^a	0.017
	(0.017)	(0.006)	(0.050)	(0.004)	(0.021)
Ln(Av. Wages)	-0.606^a	-0.021^a	-0.153^a	-0.018^a	-0.019^c
	(0.017)	(0.006)	(0.027)	(0.003)	(0.010)
Observations	126,024	126,024	126,024	126,024	126,024
F-statistic	68.10				
Mean of Dep. Var.		0.173	0.173	0.114	0.114
Ext-Margin Elasticity		0.387	-0.921	0.455	0.432

Notes: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province reported in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the log difference between the average of X in 2009-2013 and its average in 2002-2008. All specifications include firm fixed effects, province fixed effects, and sector-period fixed effects.

Table I.4: Intensive Margin: Robustness to Excluding Zip Codes Linked to Auto Industry

	Panel A: Exclude zip codes with high auto employment share		*Panel B: Exclude zip codes with at least one sizeable auto maker*		*Panel C: Exclude zip codes 'neigh-boring' zipcodes in Panel A*		*Panel D: Exclude sectors with input- -output links to automakers*	
Model:	1st Stage	2SLS	1st Stage	2SLS	1st Stage	2SLS	1st Stage	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔLn(Sales)		-2.515^a		-2.509^a		-2.478^a		-3.188^a
		(0.585)		(0.760)		(0.641)		(0.906)
ΔLn(Dist-Pop-Weighted Vehicles p.c.)	0.869^a		0.909^a		0.895^a		0.724^a	
	(0.111)		(0.157)		(0.132)		(0.127)	
ΔLn(Average Wages)	-0.490^a	-1.740^a	-0.481^a	-1.731^a	-0.469^a	-1.658^a	-0.483^a	-2.084^a
	(0.045)	(0.267)	(0.056)	(0.350)	(0.041)	(0.289)	(0.048)	(0.366)
ΔLn(TFP)	1.050^a	3.777^a	1.052^a	3.768^a	1.042^a	3.692^a	1.047^a	4.466^a
	(0.031)	(0.536)	(0.043)	(0.681)	(0.031)	(0.590)	(0.033)	(0.829)
Observations	7,180	7,180	4,595	4,595	6,131	6,131	6,072	6,072
Sector FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F-statistic	61.51		33.41		45.70		32.71	

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the log difference between the average of X in 2009-2013 and its average in 2002-2008. 'ΔLn(Dist-Pop-Wght. vehicles p.c.)' denotes the baseline instrument constructed using data on vehicles per capita at the municipal level and applying the weights from the gravity equation reported in column 1 of Table 1. 'F-statistic' denotes the corresponding statistic for the null hypothesis that the coefficient on the ΔLn(Dist-Pop weighted Vehicles p.c.) covariate is equal to zero. See text for details on the construction of each subsample. All regressions include sector fixed effects.

elasticity that enters in the construction of the instrument weights. In column 3, we use a more

Table I.5: Alternative Instruments

Dependent Variable:	ΔLn(Total Sales)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Dist-Pop-Wght. Vehicles p.c.)	0.888[a]						
Gravity: mun-mun flows (Baseline)	(0.103)						
ΔLn(Dist-Pop-Wght. Vehicles p.c.)		0.895[a]					
Baseline incl. own mun. & prov. dummies		(0.102)					
ΔLn(Dist-Pop-Wght. Vehicles p.c.)			0.632[a]				
Gravity: distance dummies			(0.070)				
ΔLn(Dist-Pop-Wght. Vehicles p.c.)				0.883[a]			
Gravity: firm-mun flows				(0.088)			
ΔLn(Weighted Vehicles p.c.)					0.257[a]		
Weights: firm-level mun. shares					(0.069)		
ΔLn(Dist-Pop-Wght. Vehicles p.c.)						0.244[a]	
Fixed coefficients: $\beta_{pop} = 1$, $\beta_{dist} = -1$						(0.070)	
ΔLn(Dist-Pop-Wght. Vehicles p.c.)							0.651[a]
Baseline in levels							(0.083)
Adão et al. (2019) std. error							(0.099)
Observations	8,009	8,009	8,009	7,906	7,850	8,009	8,009
F-statistic	75.00	77.08	80.65	99.67	13.86	12.17	61.19
							43.13
	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Total Sales)	-2.374[a]	-2.376[a]	-2.018[a]	-2.461[a]	-4.374[a]	-2.042[b]	-2.526[a]
	(0.526)	(0.521)	(0.463)	(0.447)	(1.266)	(0.841)	(0.579)
							(0.610)
Observations	8,009	8,009	8,009	7,906	7,850	8,009	8,009

Note: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered by province reported in parentheses. In column 7, standard errors computed following the procedure in Adão et al. (2019) are also reported on the following line. All specifications include firm-level log TFP and log average wages as additional controls (coefficients not included to save space). Additionally, all specifications also include sector fixed effects.

flexible specification with dummies for distance intervals; we use the estimated coefficients on these distance intervals to construct the instrument weights. In column 4, we use data on firm-to-municipality flows to estimate the elasticity of trade flows to distance. In column 5, we construct our instrument equating the instrument weights corresponding to each firm to the 2006 observed municipality-specific domestic sales shares of the corresponding firm. In column 6, we impose fixed coefficients on the gravity equation, equal to 1 for population and -1 for distance. Finally, in column 7, we use an instrumental variable that is identical to our baseline instrument expect for the fact that we use as instrument the boom-to-bust change in the level (instead of the log) of the weighted average of the number of vehicles per capita in each of the two periods. The instrument reported in column 7 is thus a specific case of the type of shift-share instrument discussed recently by Adão et al. (2019), Borusyak et al. (2020) and Goldsmith-Pinkham et al. (2020), among others. Building on this literature, we report standard errors for the first-stage and for the second-stage coefficients that we compute following the procedure described in Adão et al. (2019). The standard errors estimated with this procedure are only marginally larger than the ones clustered by province, and therefore do not affect the significance of the results. The second-stage estimates of the elasticity of exports with respect to total sales are all broadly similar to those in the baseline (i.e., the estimated elasticities in columns 2 to 7 are very similar to the elasticity reported in column 1). The only exception is the estimated reported in column 5, which is substantially larger in absolute value (-4.374) than our baseline estimate.

In Table I.6, we assess whether our estimates are affected by the inclusion of potential confounders, as in Table 8 in the paper. The elasticity of the boom-to-bust change in log exports with respect to the change in log total sales is stable around -2.4 in all specifications.

Next, we check in Table I.7 whether the estimates with total sales as the key right-hand-side variable are sensitive to using an alternative measure of TFP based on value added (instead of total sales, as in the baseline specification), replicating Table 9 in the paper. In this case, the estimated elasticity is -1.450, somewhat smaller in absolute value than the baseline result but still statistically significant at any commonly used significance level.

Table I.6: Confounding Factors

Dependent Variable:	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Total Sales)	-2.374^a	-2.481^a	-2.395^a	-2.425^a	-2.469^a	-2.522^a	-2.383^a
	(0.526)	(0.549)	(0.560)	(0.556)	(0.527)	(0.561)	(0.560)
ΔShare of Temp. Workers		-0.404^a					
(firm level)		(0.135)					
ΔShare of Temp. Workers			-0.094				
(munic. level)			(0.188)				
ΔManufacturing Empl. p.c.				0.354^a			
(munic. level)				(0.068)			
ΔLn(Financial Costs)					-0.061^a		
(firm level)					(0.019)		
Financial Costs in Boom						-0.014	
(firm level)						(0.019)	
Financial Costs in Boom							-0.063
(munic. level)							(0.047)
F-Statistic	75.00	74.48	73.55	73.51	71.23	64.96	73.28

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province reported in parentheses. All specifications include firm-level log TFP and log wages as additional controls (coefficients not included to save space). All specifications also include sector fixed effects.

Finally, we close this section by exploring whether the increase in exports in reaction to a common demand-driven drop in domestic sales is indeed larger for those firms whose short-run marginal cost function is steeper or, equivalently, for those firms whose elasticity of output with respect to flexible inputs is lower (we rely on our production function estimates in Appendix F.2 to measure these output elasticities.) The results are in Table I.8. Notice from the table that the elasticity of exports with respect to total sales is lower in sectors with higher elasticities with respect to materials (columns 1 and 4), in sectors with a higher elasticity with respect to labor (columns 2 and 5), and in sectors with a higher elasticity of output with respect to the use of temporary workers (columns 3 and 6). Notice, however, that only the two latter results are statistically significant at standard levels, and only for our gravity-based estimate. Interestingly, the estimates in columns 3 and 6 imply that we cannot rule out that the elasticity of exports with respect to a domestic demand-driven change in total sales equals 0 for firms that satisfy two conditions: (a) their elasticity of output with respect to labor equals 1; (b) their share of temporary workers in their total workforce also equals 1. This prediction is consistent with the model described at the beginning of this section and the micro-foundation in Appendix A, as these firms would have short-run constant marginal costs according to this micro-foundation.

Table I.7: Alternative TFP Measures

Dependent Variable:	ΔLn(Exports)			
	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
ΔLn(Total Sales)	0.724[a]	-2.374[a]	0.850[a]	-1.450[a]
	(0.050)	(0.526)	(0.046)	(0.339)
ΔLn(Average Wages)	-0.217[a]	-1.750[a]	-0.514[a]	-1.174[a]
	(0.063)	(0.250)	(0.074)	(0.132)
ΔLn(TFP Sales): Baseline	0.509[a]	3.690[a]		
	(0.055)	(0.482)		
ΔLn(TFP Value-Added)			0.685[a]	1.590[a]
			(0.066)	(0.157)
Observations	8,009	8,009	8,009	8,009
F-Statistic		75.00		80.74

Note: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered at the province level are reported in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the difference in $\text{Ln}(X)$ between its average in the 2009-2013 period and its average in the 2002-2008 period. All specifications include sector fixed effects.

Table I.8: Heterogeneous Effects with Total Sales: Second Stage

Dependent Variable:	ΔLn(Exports)					
Instrument for ΔLn(Total Sales):	Municipality-level IV			Gravity-Based IV		
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Total Sales)	-2.750[b]	-2.671[a]	-2.194[a]	-2.508[a]	-3.635[a]	-2.688[a]
	(1.165)	(0.979)	(0.793)	(0.755)	(1.134)	(0.579)
ΔLn(Total Sales) × High Output elasticity wrt Materials	1.434			0.462		
	(1.358)			(0.782)		
ΔLn(Total Sales) × High Output elasticity wrt Labor		1.335			2.205[c]	
		(1.263)			(1.143)	
ΔLn(Total Sales) × (High Output elast. wrt Labor × Temp. Ratio			3.037			3.963[a]
			(2.364)			(1.350)
Observations	8,009	8,009	7,889	8,009	8,009	7,889
Sector FE	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	No	No	No
P-value for $H_0 : \beta_1 + \beta_2 = 0$	0.06	0.08	0.65	0.00	0.00	0.23

Note: [a] denotes 1% significance, [b] denotes 5% significance, [c] denotes 10% significance. Standard errors clustered at the municipality level (in columns 1 to 3) and province level (in columns 4 to 6) reported in parenthesis. For any X, $\Delta\text{Ln}(X)$ is the difference in $\text{Ln}(X)$ between its average in the 2009-13 period and its average in the 2002-08 period. The output elasticities with respect to inputs are estimated with the same production function we use to estimate TFP.

J Details on Counterfactual Analysis

J.1 System of Equations for Counterfactual Exercise

We describe here the step-by-step derivation of the three types of equations we use in the counterfactual analysis described in section VII.

The first two equation types map the boom-to-bust counterfactual changes in exports and domestic sales of each firm to the counterfactual changes in the aggregate demand shifter and price index in the firm's sector. The step-by-step derivations of these first two equation types are analogous and, thus, we show these derivations in parallel.

By defining $B_{sx} \equiv E_{sx}/P_{sx}$ and $B_{sd} \equiv E_{sd}/P_{sd}$, and combining equation (16) with both equation (15) and its analogous expression for domestic sales, we can write the log boom-to-bust change in exports and domestic sales of a firm i as

$$\ln\left[\frac{R_{ix1}}{R_{ix0}}\right] = \ln\left[\frac{B_{sx1}}{B_{sx0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right] + \sigma\ln\left[\frac{P_{sx1}}{P_{sx0}}\right] - \frac{(\sigma-1)\,\lambda}{1+\lambda}\ln\left[\frac{R_{i1}}{R_{i0}}\right], \tag{J.1a}$$

$$\ln\left[\frac{R_{id1}}{R_{id0}}\right] = \ln\left[\frac{B_{sd1}}{B_{sd0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right] + \sigma\ln\left[\frac{P_{sd1}}{P_{sd0}}\right] - \frac{(\sigma-1)\,\lambda}{1+\lambda v}\ln\left[\frac{R_{i1}}{R_{i0}}\right], \tag{J.1b}$$

where $\ln[x_1/x_0]$ denotes the log change between the boom and the bust periods in any covariate x, and remember that

For any variable x, we define as x'_1 the counterfactual value that this variable takes in the bust period in our quantification if the demand shifters take in the bust period the counterfactual values $\{B'_{sd1}\}_{s=1}^{S}$, and all other demand and supply shocks change between boom and bust periods as they actually did. Therefore, analogously to equations (J.1a) and (J.1b), we can define the following two equations

$$\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] = \ln\left[\frac{B_{sx1}}{B_{sx0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right] + \sigma\ln\left[\frac{P_{sx1}}{P_{sx0}}\right] - \frac{(\sigma-1)\,\lambda}{1+\lambda}\ln\left[\frac{R'_{i1}}{R_{i0}}\right] \tag{J.2a}$$

$$\ln\left[\frac{R'_{id1}}{R_{id0}}\right] = \ln\left[\frac{B'_{sd1}}{B_{sd0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right] + \sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\right] - \frac{(\sigma-1)\,\lambda}{1+\lambda}\ln\left[\frac{R'_{i1}}{R_{i0}}\right]. \tag{J.2b}$$

Equations (J.2a) and (J.2b) allow us to compute the impact of counterfactual demand shocks $\ln[B'_{sd1}/B_{sd0}]$ on firms' domestic sales and exports while holding the changes in the foreign price index, P_x, and in the equilibrium wages that each firm i faces, ω_i, unaltered by the counterfactual change in the demand shocks. In the case of non-exporting firms, only equation (J.2b) applies for all these firms. Equations (J.2a) and (J.2b) illustrate that the counterfactual changes in exports and domestic sales of a firm i that belongs to a sector s are a function of the actual changes in firm

i's own supply and idiosyncratic demand shifters

$$\left\{ \ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right], \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right], \ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right], \ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right], \ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right], \ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right] \right\} \tag{J.3}$$

and, through the counterfactual change in the domestic price index, P'_{sd1}/P_{sd0}, of the actual changes in the supply and idiosyncratic demand shifters of all other firms in the same sector s. The variables listed in equation (J.3) are unobserved in our data. However, combining equations (J.1a) and (J.1b) with equations (J.2a) and (J.2b), respectively, we can rewrite equation (J.2a) as

$$\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] = \ln\left[\frac{R_{ix1}}{R_{ix0}}\right] - \frac{(\sigma-1)\,\lambda}{1+\lambda}\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \tag{J.4}$$

where $\chi_0 \equiv R_{ix0}/(R_{id0}+R_{ix0})$ denotes the initial export share of firm i, and

$$\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0}) \qquad \text{and} \qquad \frac{R_{i1}}{R_{i0}}$$

denote, respectively, the counterfactual and observed change in firm i's total sales. Equation (J.4) describes the first type of equations we use in our counterfactual analysis; we use one such equation for each firm with positive exports in both the boom and the bust.

Similarly to how we obtained the expression in equation (J.4), we can rewrite equation (J.2b) as

$$\begin{aligned}\ln\left[\frac{R'_{id1}}{R_{id0}}\right] = {} & \ln\left[\frac{B'_{sd1}}{B_{sd0}}\left(\frac{B_{sd1}}{B_{sd0}}\right)^{-1}\right] + \sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right] + \ln\left[\frac{R_{id1}}{R_{id0}}\right] \\ & - \frac{(\sigma-1)\,\lambda}{1+\lambda}\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right],\end{aligned} \tag{J.5}$$

where

$$\frac{B'_{sd1}}{B_{sd0}}\left(\frac{B_{sd1}}{B_{sd0}}\right)^{-1} \qquad \text{and} \qquad \frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1},$$

denote the counterfactual change (relative to the actual change) in the aggregate sectoral demand shifter and price index, respectively. Equation (J.5) describes the second type of equations we use in our counterfactual analysis; we use one such equation for each firm with positive domestic sales in both the boom and the bust.

Besides equations (J.4) and (J.5), the system of equations we use to perform our quantification includes equations of a third type that endogenize the counterfactual change in the sector-specific domestic price indices. To derive these equations (one for each sector), it is useful to write the sectoral domestic price index in any sector s and period t as

$$P_{sdt} = \frac{E_{sdt}}{B_{sdt}} = \frac{R_{sdt} + R^{\mathcal{X}}_{sdt}}{B_{sdt}}, \tag{J.6}$$

where R_{sdt} denotes the aggregate domestic sales of firms located in country d and operating in sector s, and $R^{\mathcal{X}}_{sdt}$ denotes the aggregate imports of country d in sector s (i.e., total sales in country d by all firms located in the foreign country). We can thus write the relative change in the domestic price index between the boom and bust periods in sector s as

$$\frac{P_{sd1}}{P_{sd0}} = \frac{R_{sd1} + R^{\mathcal{X}}_{sd1}}{R_{sd0} + R^{\mathcal{X}}_{sd0}} \frac{B_{sd0}}{B_{sd1}}$$

or, equivalently,

$$\frac{P_{sd1}}{P_{sd0}} = \left(\frac{R_{sd0}}{R_{sd0} + R^{\mathcal{X}}_{sd0}} \frac{R_{sd1}}{R_{sd0}} + \frac{R^{\mathcal{X}}_{sd0}}{R_{sd0} + R^{\mathcal{X}}_{sd0}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} \right) \frac{B_{sd0}}{B_{sd1}}.$$

Simplifying notation, we can write that

$$\frac{P_{sd1}}{P_{sd0}} = \left(s^{\mathcal{D}}_{sd0} \frac{R_{sd1}}{R_{sd0}} + (1 - s^{\mathcal{D}}_{sd0}) \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} \right) \left(\frac{B_{sd1}}{B_{sd0}} \right)^{-1},$$

where $s^{\mathcal{D}}_{sd0}$ is the boom-period share of total consumption in country d spent in varieties produced by firms located in the same country d. Noting that

$$\frac{R_{sd1}}{R_{sd0}} = \sum_{i \in \mathcal{D}_s} s^{\mathcal{D}}_{id0} \frac{R_{id1}}{R_{id0}},$$

we can rewrite the log counterfactual change in the price index P_{sd} relative to the actual change as

$$\begin{aligned} \ln \left[\frac{P'_{sd1}}{P_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}} \right)^{-1} \right] &= \ln \left(s^{\mathcal{D}}_{sd0} \sum_{i \in \mathcal{D}_s} s^{\mathcal{D}}_{id0} \frac{R'_{id1}}{R_{id0}} + (1 - s^{\mathcal{D}}_{sd0}) \frac{(R^{\mathcal{X}}_{sd1})'}{R^{\mathcal{X}}_{sd0}} \right) \\ &\quad - \ln \left(s^{\mathcal{D}}_{sd0} \sum_{i \in \mathcal{D}_s} s^{\mathcal{D}}_{id0} \frac{R_{id1}}{R_{id0}} + (1 - s^{\mathcal{D}}_{sd0}) \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} \right) - \ln \left[\frac{B'_{sd1}}{B_{sd0}} \left(\frac{B_{sd1}}{B_{sd0}} \right)^{-1} \right]. \end{aligned} \tag{J.7}$$

A key element in this expression is the variable $(R^{\mathcal{X}}_{sd1})'/R^{\mathcal{X}}_{sd0}$, which denotes the counterfactual total change in imports to country d in sector s; i.e., counterfactual change in Spanish imports in sector s. Without loss of generality, we can rewrite

$$\begin{aligned} \frac{(R^{\mathcal{X}}_{sd1})'}{R^{\mathcal{X}}_{sd0}} &= \frac{(R^{\mathcal{X}}_{sd1})'}{R^{\mathcal{X}}_{sd0}} \left[\frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} \right]^{-1} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} = \frac{\frac{\sum_{i \in \mathcal{X}_s} R'_{id1}}{\sum_{i \in \mathcal{X}_s} R_{id0}}}{\frac{\sum_{i \in \mathcal{X}_s} R_{id1}}{\sum_{i \in \mathcal{X}_s} R_{id0}}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} = \frac{\sum_{i \in \mathcal{X}_s} \left(\frac{R_{id0}}{\sum_{i \in \mathcal{X}_s} R_{id0}} \right) \frac{R'_{id1}}{R_{id0}}}{\sum_{i \in \mathcal{X}_s} \left(\frac{R_{id0}}{\sum_{i \in \mathcal{X}_s} R_{id0}} \right) \frac{R_{id1}}{R_{id0}}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} \\ &= \frac{\sum_{i \in \mathcal{X}_s} s^{\mathcal{X}}_{id0} \frac{R'_{id1}}{R_{id0}}}{\sum_{i \in \mathcal{X}_s} s^{\mathcal{X}}_{id0} \frac{R_{id1}}{R_{id0}}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}} = \frac{\sum_{i \in \mathcal{X}_s} s^{\mathcal{X}}_{id0} \frac{P'_{id1} Q'_{id1}}{P_{id0} Q_{id0}}}{\sum_{i \in \mathcal{X}_s} s^{\mathcal{X}}_{id0} \frac{P_{id1} Q_{id1}}{P_{id0} Q_{id0}}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}, \end{aligned}$$

where $s^{\mathcal{X}}_{id0}$ is the share of firm i in total sales in market d by firms located in x (i.e., by firms belonging to the set $\mathcal{X}$); i.e., share of total imports in market d that correspond to firm i.

In general, P'_{id1} will differ from P_{id1}; i.e., differences in the aggregate demand shock in country d affect the total quantity produced of all the firms located in country x and, thus, affect their marginal cost and prices. However, assuming that market d is small for the firms located in country x (i.e., only a very small share of total sales of firms located in country x correspond to sales in country d; country d is "small" for foreign firms), it will be true that

$$P'_{id1} = P_{id1},$$

for all firms located in country x. Therefore, we can simplify the expression for the counterfactual change in Spanish imports in sector s as

$$\frac{(R^{\mathcal{X}}_{sd1})'}{R^{\mathcal{X}}_{sd0}} = \frac{\sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \frac{P_{id1}}{P_{id0}} \frac{Q'_{id1}}{Q_{id0}}}{\sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \frac{P_{id1}}{P_{id0}} \frac{Q_{id1}}{Q_{id0}}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}, \tag{J.8}$$

and we can write

$$\frac{Q'_{id1}}{Q_{id0}} = \left(\frac{P_{id1}}{P_{id0}}\right)^{-\sigma} \frac{B'_{sd1}}{B_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}, \tag{J.9}$$

$$\frac{Q_{id1}}{Q_{id0}} = \left(\frac{P_{id1}}{P_{id0}}\right)^{-\sigma} \frac{B_{sd1}}{B_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}, \tag{J.10}$$

where, as we have previously done for the case of the firms located in Spain, we set the change in the idiosyncratic demand shocks of the foreign firms to equal the actual change (i.e., $\xi'_{id1} = \xi_{id1}$) with the aim of having a counterfactual that isolates the impact of the aggregate domestic demand shock. Therefore, plugging equations (J.9) and (J.10) into equation (J.8), we can further rewrite the expression for the counterfactual change in Spanish imports in sector s as

$$\begin{aligned}
\frac{(R^{\mathcal{X}}_{sd1})'}{R^{\mathcal{X}}_{sd0}} &= \frac{\sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \frac{B'_{sd1}}{B_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}}{\sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \frac{B_{sd1}}{B_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}, \\
&= \frac{\frac{B'_{sd1}}{B_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma} \sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}}{\frac{B_{sd1}}{B_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma} \sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}, \\
&= \frac{\frac{B'_{sd1}}{B_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma}}{\frac{B_{sd1}}{B_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}.
\end{aligned}$$

Plugging this expression back into equation (J.7), we obtain:

$$\begin{aligned}
\ln\left[\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right] = {} & \ln\left(s^{\mathcal{D}}_{sd0} \sum_{i\in\mathcal{D}_s} s^{\mathcal{D}}_{id0} \frac{R'_{id1}}{R_{id0}} + (1 - s^{\mathcal{D}}_{sd0}) \frac{B'_{sd1}}{B_{sd0}} \left(\frac{B_{sd1}}{B_{sd0}}\right)^{-1} \left(\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right)^{\sigma} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}\right) \\
& - \ln\left(s^{\mathcal{D}}_{sd0} \sum_{i\in\mathcal{D}_s} s^{\mathcal{D}}_{id0} \frac{R_{id1}}{R_{id0}} + (1 - s^{\mathcal{D}}_{sd0}) \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}\right) - \ln\left[\frac{B'_{sd1}}{B_{sd0}}\left(\frac{B_{sd1}}{B_{sd0}}\right)^{-1}\right].
\end{aligned} \tag{J.11}$$

This is the third equation type we use to compute the predictions of our counterfactual analysis; we use one such equation for each sector.

Summing up, the system of equations we use to compute the predictions of our counterfactual analysis uses an equation like that in equation (J.4) for every firm exporting in both the boom and the bust, an equation like that in equation (J.5) for every firm with positive domestic sales in both the boom and the bust, and one equation like that in equation (J.11) for every sector.

J.2 Decomposition of the Variance of Boom-to-Bust Changes in Total Sales

We can rewrite equation (17) as

$$\Delta \ln \mathcal{R}_{ix} = \beta \Delta \ln \mathcal{R}_i + \varepsilon_{ix}, \tag{J.12}$$

with

$$\beta = -\frac{(\sigma - 1)\,\lambda}{1+\lambda} \qquad \text{and} \qquad \varepsilon_{ix} \equiv u^{\xi}_{ix} + \frac{(\sigma-1)}{1+\lambda}(u^{\varphi}_i - u^{\omega}_i), \tag{J.13}$$

where, as in equation (11) in the main text, we denote by $\Delta \ln \mathcal{X}$ the residual of a regression of a variable $\Delta \ln X$ on a set of sector fixed effects $\{d\}_s$, location fixed effects $\{d\}_\ell$, and the observable covariates $\Delta \ln \varphi^*_i$, and $\Delta \ln \omega^*_i$. Using this notation, we can write the probability limit of the OLS and IV estimators of β as

$$\beta_{ols} = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}_i)}{var(\Delta \ln \mathcal{R}_i)}, \qquad\qquad \beta_{iv} = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}^*_i)}{cov(\Delta \ln \mathcal{R}_i, \Delta \ln \mathcal{R}^*_i)}, \tag{J.14}$$

where $\Delta \ln \mathcal{R}^*_i$ is the part of $\Delta \ln \mathcal{R}_i$ that is mean-independent of the residual of the structural equation, ε_{ix}; i.e., $\Delta \ln \mathcal{R}^*_i \equiv \Delta \ln \mathcal{R}_i - \mathbb{E}[\Delta \ln \mathcal{R}_i | \varepsilon_{ix}]$. Denoting $\Delta \ln \mathcal{R}^{\varepsilon}_i = \mathbb{E}[\Delta \ln \mathcal{R}_i|\varepsilon_{ix}]$, we can thus rewrite $\Delta \ln \mathcal{R}_i = \Delta \ln \mathcal{R}^*_i + \Delta \ln \mathcal{R}^{\varepsilon}_i$. In practice, given an estimate $\hat{\beta}_{iv}$ of β_{iv}, we recover an estimate of ε_{ix} for every exporter i as $\Delta \ln \mathcal{R}_{ix} - \hat{\beta}_{iv}\Delta \ln \mathcal{R}_i$; i.e., $\hat{\varepsilon}_{ix} \equiv \Delta \ln \mathcal{R}_{ix} - \hat{\beta}_{iv}\Delta \ln \mathcal{R}_i$. Given this estimate $\hat{\varepsilon}_{ix}$, we compute an estimate of $\Delta \ln \mathcal{R}^{\varepsilon}_i$ by running a regression of $\Delta \ln \mathcal{R}_i$ on $\hat{\varepsilon}_{ix}$ and equating our estimate of $\Delta \ln \mathcal{R}^{\varepsilon}_i$ to the predicted value of such regression.

Given the expressions for β_{ols} and β_{iv} in equation (J.14), after simple algebraic manipulations, we can relate β_{ols} and β_{iv} as

$$\beta_{ols} = \beta_{iv}\frac{var(\Delta \ln \mathcal{R}^*_i)}{var(\Delta \ln \mathcal{R}_i)} + \beta_{\varepsilon}\left(1 - \frac{var(\Delta \ln \mathcal{R}^*_i)}{var(\Delta \ln \mathcal{R}_i)}\right), \tag{J.15}$$

where

$$\beta_{\varepsilon} = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}^{\varepsilon}_i)}{cov(\Delta \ln \mathcal{R}_i, \Delta \ln \mathcal{R}^{\varepsilon}_i)} = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}^{\varepsilon}_i)}{var(\Delta \ln \mathcal{R}^{\varepsilon}_i)}. \tag{J.16}$$

Given equation J.15, we can compute the share of the variance in total sales that is due to factors mean-independent of ε_{ix} as

$$\frac{var(\Delta \ln \mathcal{R}^*_i)}{var(\Delta \ln \mathcal{R}_i)} = \frac{\beta_{ols} - \beta_{\varepsilon}}{\beta_{iv} - \beta_{\varepsilon}}. \tag{J.17}$$

Given consistent estimates of β_{ols}, β_{iv} and β_{ε}, we use this expression to compute a consistent estimate of $var(\Delta \ln \mathcal{R}^*_i)/var(\Delta \ln \mathcal{R}_i)$. When performing this calculation using our observed data, we obtain that this ratio of variances is equal to 35%. Given the definition of ε_x in equation (J.13), we can thus conclude that 35% of the variance of the residual of projecting the boom-to-bust log change in firms' total sales on sector and location fixed effects and the observable covariates $\Delta \ln \varphi^*_i$ and $\Delta \ln \omega^*_i$ is due to factors that are mean-independent of the unobserved supply shocks u^{φ}_i and u^{ω}_i and the export demand shocks u^{ξ}_{ix}.

We also perform a similar analysis to that described in equations (J.12) to (J.17) but with the aim of decomposing the cross-firm variance in the observed boom-to-bust log changes in total sales, $var(\Delta \ln R_i)$, into a component that is the result of projecting these changes on the regression residual

$$\tilde{\varepsilon}_{ix} = \Delta \ln R_{ix} - \beta \Delta \ln R_i = \gamma_{sx} + \gamma_{\ell x} + \frac{(\sigma - 1)}{1 + \lambda} \delta_{\varphi} \Delta \ln \varphi_i^* - \frac{(\sigma - 1)}{1 + \lambda} \delta_{\omega} \Delta \ln \omega_i^* + \varepsilon_{ix},$$

and a component that is due to the impact on $\Delta \ln(R_i)$ of variables that are mean-independent of $\tilde{\varepsilon}_{ix}$. When performing this variance decomposition, we find that the variables orthogonal to $\tilde{\varepsilon}_{ix}$ explain 41% of the variance in the observed changes in total sales; i.e., $var(\Delta \ln R_i^*)/var(\Delta \ln R_i) = 0.41$. It is important to remark that this alternative variance decomposition requires assuming that our instrument is valid unconditionally, and not just conditionally on sector and location fixed effects and on our proxies for firms' factor prices and productivity.

J.3 Counterfactual Exercise With Boom-to-Bust Changes in Trade Costs

We show here results from a quantification of how much more the total sales of Spanish firms would have dropped if firms had faced an increase in export costs in the bust, when trying to substitute domestic markets for export markets.

To quantify the sensitivity of the boom-to-bust change in total sales to the role played by export markets as facilitators of the venting out of products whose demand in Spain dropped between the boom and bust, we compute the counterfactual growth in exports that we would have observed if, simultaneously with the change in the aggregate domestic demand shifters $\{B_{sd}\}_{s=1}^{S}$, we had observed a change in the export trade costs $\{\tau_{sx}\}_{s=1}^{S}$ between the boom and the bust periods. To compute such counterfactual, we use a variant of the system defined by equations (J.4), (J.5) and (J.11), all of them described in Appendix J.1. Specifically, our counterfactual analysis relies on equation (J.5), equation (J.11), and the following equation

$$\begin{aligned} \ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] &= \ln\left[\frac{R_{ix1}}{R_{ix0}}\right] - (\sigma - 1)\ln\left[\frac{\tau'_{sx1}}{\tau_{sx0}}\left(\frac{\tau_{sx1}}{\tau_{sx0}}\right)^{-1}\right] \\ &\quad - \frac{(\sigma - 1)\lambda}{1 + \lambda}\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1 - \chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \end{aligned} \tag{J.18}$$

which is a generalization of the expression in equation (J.4). We focus on counterfactual exercises in which the counterfactual relative change in trade costs is constant across sectors; i.e.,

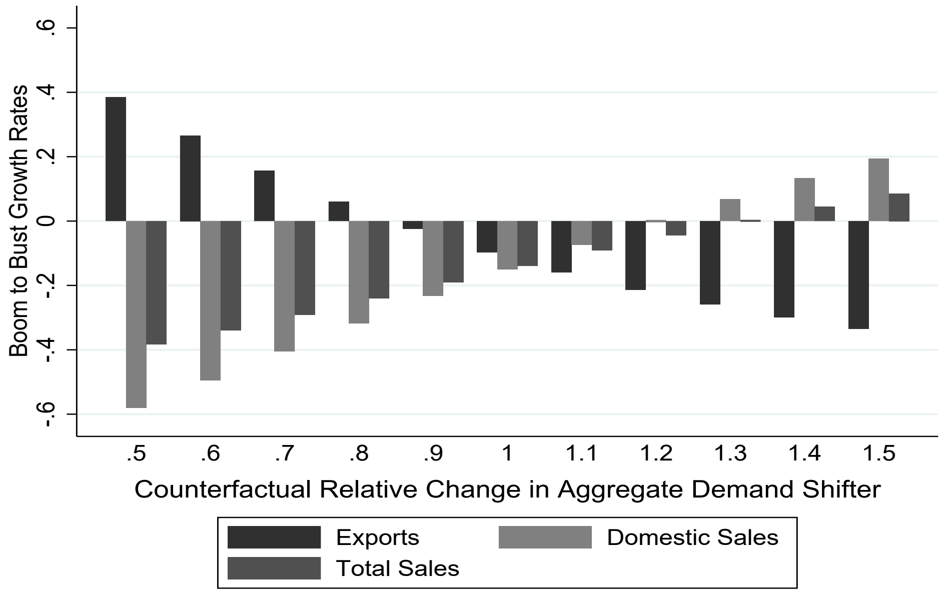
$$\frac{\tau'_{sx1}}{\tau_{sx0}}\left(\frac{\tau_{sx1}}{\tau_{sx0}}\right)^{-1} = \frac{\tau'_{sx1}}{\tau_{sx1}} = \Gamma_{\tau}. \tag{J.19}$$

Panels (a) and (b) in Figure J.1 illustrate the results of our counterfactual analysis for $\Gamma_\tau = 1.10$ and for $\Gamma_\tau = 1.25$, respectively. As a comparison between the two panels featured in Figure J.1, and between either of them and that in Figure 4, illustrates, a boom-to-bust change in trade costs would have severely impacted the growth in exports between the boom and the bust. For any value of Γ_B, a growth in trade costs of only 10% already has a quantitatively important negative effect on aggregate exports and, although domestic sales react to this change in trade costs by increasing by more (or, more precisely, decreasing by less) than they would otherwise have, the overall impact of the boom-to-bust change in trade costs on the change in total sales is negative.

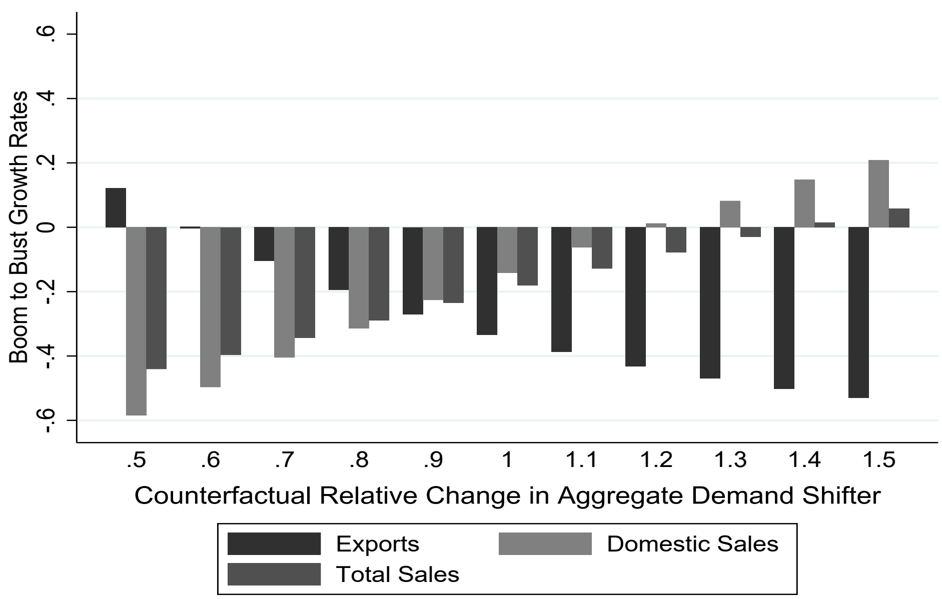
As an example, at the value of Γ_B that maintains aggregate demand shifters constant between the boom and the bust periods (i.e., $\Gamma_B = 1.09$), if no change in trade costs had taken place, aggregate exports would have grown by 5.79%, domestic sales would have fallen by 9.1% and total sales would have fallen by 6.04% (see section VII). Conversely, if trade costs had increased in 10% more than they actually did between the boom and the bust periods (i.e., $\Gamma_\tau = 1.10$), then the

Figure J.1: Impact of Aggregate Demand Shocks With Simultaneous Changes in Trade Costs

(a) With a 10% Increase in Trade Costs ($\Gamma_\tau = 1.1$)



(b) With a 25% Increase in Trade Costs ($\Gamma_\tau = 1.25$)



Notes: The horizontal axis indicates the value of Γ_B. The export and domestic sales growth rates indicated in the vertical axis correspond to those predicted by equations (J.5), (J.11), (J.18) and (J.19). Panel (a) imposes $\Gamma_\tau = 1.1$; Panel (b) imposes $\Gamma_\tau = 1.25$. Given these counterfactual growth rates in export and domestic sales, we compute the counterfactual growth rate in total sales as $(R'_{ix1}/R_{ix0})\chi_{i0} + (R'_{id1}/R_{id0})(1 - \chi_{i0})$.

boom-to-bust growth in aggregate exports would have actually been negative and equal to -15.37%, domestic sales would have fallen by 'only' 8.07% (i.e., less than in the baseline with no change in trade costs) and total sales would have fallen by 9.56%. If the growth in trade costs had been 25%, then the drop in exports, domestic sales and total sales in the counterfactual scenario in which aggregate demand shifters had remained constant between the boom and bust periods would have been 38.25%, 7.01% and 13.37%, respectively. These numbers illustrate that aggregate exports are very sensitive to changes in trade costs and, although domestic sales partly react to these changes in trade costs by compensating for the fall in exports (because of the substitutability between domestic and foreign markets), aggregate total sales are still quite sensitive to trade costs.

If we were to evaluate the impact of a change in trade costs while holding the boom-to-bust change in the aggregate demand shifters at their actual value (i.e., $\Gamma_B = 1$), our model predicts that the drop in total sales would have been 10.23% if trade costs had changed as they did in the data, 13.97% if trade costs in the bust were 10% larger, and 18.11% if they were 25% larger. Importantly, these changes in total sales partly reflect the impact that the change in trade costs has on domestic sales according to our model; specifically, our model predicts that we would have observed a drop in domestic sales of 15.91% if trade costs had changed as they did in the data, a drop of 15.05% if trade costs in the bust were 10% larger, and a drop of 14.2% if trade costs were 25% larger.

J.4 Counterfactual Exercise Under Alternative Parameter Values

We explore here how robust the results of our baseline quantification are to different values of the parameter $(\sigma - 1)\lambda/(1 + \lambda)$.

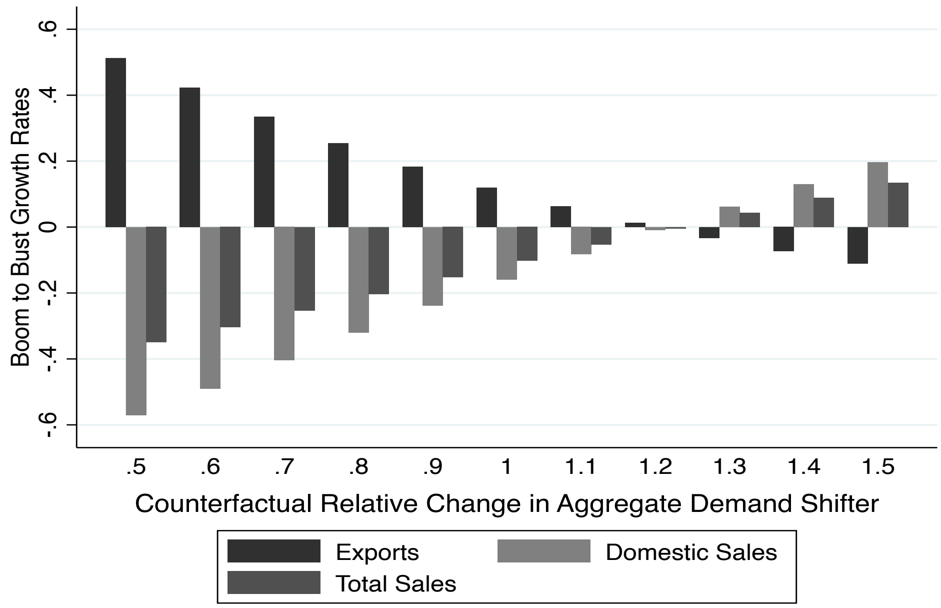
To quantify the impact that the value of the parameter $(\sigma - 1)\lambda/(1 + \lambda)$ has on our results, we recompute the baseline counterfactual exercise for several different values of this parameter.

Panels (a) and (b) in Figure J.2 present results for a value of $(\sigma-1)\lambda/(1+\lambda)$ that is 25% smaller than the baseline (i.e., $(\sigma - 1)\lambda/(1 + \lambda) = 0.75 \times 2.374 = 1.7805$) and for a value of this parameter that is 25% larger than the baseline (i.e., $(\sigma - 1)\lambda)(1 + \lambda) = 1.25 \times 2.374 = 2.9675$). The results in panel (a) show that, if aggregate demand shifters had remained invariant between the boom and the bust ($\Gamma_B = 1.09$), export sales would have increased by 6.86%, aggregate domestic sales would have dropped by 9.02%, and total sales would have dropped by 5.78%; thus, in this case, we would conclude that the vent-for-surplus mechanism explains $(11.99\% - 6.86\%)/11.99\% = 42.79\%$ of the total growth in exports. Conversely, according to the results in panel (b) (i.e., for a relatively high elasticity of exports to total sales), if aggregate demand shifters had remained invariant between the boom and the bust ($\Gamma_B = 1.09$), our model predicts that export sales would have increased by 4.87%, aggregate domestic sales would have dropped by 9.16%, and total sales would have dropped by 6.3%; thus, in this case, we would conclude that the vent-for-surplus mechanism explains $(11.99\% - 4.87\%)/11.99\% = 59.38\%$ of the total growth in exports.

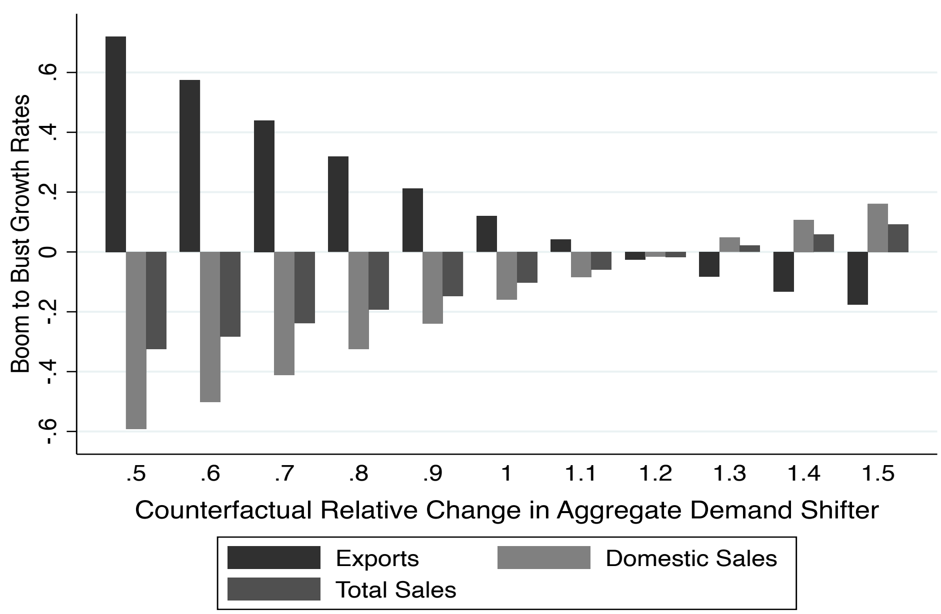
In Table J.1, we present results on the implied contribution of the vent-for-surplus mechanism to the boom-to-bust growth in exports for several additional values of the parameter $(\sigma - 1)\lambda/(1+\lambda)$. As the results illustrate, increases in the value of $(\sigma - 1)\lambda/(1 + \lambda)$ increase the implied contribution of the vent-for-surplus mechanism. However, the effect is non-linear: constant increases in $(\sigma - 1)\lambda/(1 + \lambda)$ have a decreasing effect. E.g., increasing the value of $(\sigma - 1)\lambda/(1 + \lambda)$ from 1.1870 to 1.7805 ($1.7805 - 1.1870 = 0.5935$) has an effect on the vent-for-surplus contribution of approximately 11 percentage points ($42.79\% - 31.94\%$); increasing its value from 4.1545 to 4.7480 (whose difference also equals 0.5935) has an effect of approximately 5 percentage points ($76.98\% - 71.81\%$).

Figure J.2: Impact of Aggregate Demand Shocks For Different Elasticities

(a) Low Elasticity of Exports With Respect to Total Sales ($(\sigma - 1)\lambda/(1 + \lambda) = 1.78$)



(b) High Elasticity of Exports With Respect to Total Sales ($(\sigma - 1)\lambda/(1 + \lambda) = 2.97$)



Notes: The horizontal axis indicates the value of Γ_B. The export and domestic sales growth rates indicated in the vertical axis correspond to those predicted by equations (J.4), (J.5) and (J.11). Given these counterfactual growth rates in export and domestic sales, we compute the counterfactual growth rate in total sales as $(R'_{ix1}/R_{ix0})\chi_{i0} + (R'_{id1}/R_{id0})(1 - \chi_{i0})$.

Table J.1: Sensitivity to Different Values of $(\sigma-1)\lambda/(1+\lambda)$

Value of $(\sigma-1)\lambda/(1+\lambda)$	Vent-for-surplus contribution
$0.50 \times 2.374 = 1.1870$	31.97%
$0.75 \times 2.374 = 1.7805$	42.77%
2.374	51.74%
$1.25 \times 2.374 = 2.9675$	59.38%
$1.50 \times 2.374 = 3.5610$	66.02%
$2 \times 2.374 = 4.7480$	77.01%

J.5 Counterfactual Exercise With Firm-Specific Upward-Sloping Labor Supply

We quantify here how our baseline quantification would change if we were to allow the wages that firms' face to change as they move along their marginal cost curves. To compute such counterfactual, we use a variant of the system defined by equations (J.4), (J.5) and (J.11). Specifically, while equation (J.11) is unaffected by the possible changes in firms' wages, equations (J.4) and (J.5) need to be re-derived to account for such potential changes in wage levels.

As a first-step towards re-deriving equations (J.4) and (J.5), we define two equations that generalize equations (J.2a) and (J.2b) in that they allow the equilibrium wages that each firm i faces in the bust period, ω_{i1}, to vary in the counterfactual:

$$\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] = \ln\left[\frac{B_{sx1}}{B_{sx0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega'_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right], + \sigma\ln\left[\frac{P_{sx1}}{P_{sx0}}\right] - \frac{(\sigma-1)\,\lambda}{1+\lambda}\ln\left[\frac{R'_{i1}}{R_{i0}}\right] \tag{J.20a}$$

$$\ln\left[\frac{R'_{id1}}{R_{id0}}\right] = \ln\left[\frac{B'_{sd1}}{B_{sd0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega'_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right] + \sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\right] - \frac{(\sigma-1)\,\lambda}{1+\lambda}\ln\left[\frac{R'_{i1}}{R_{i0}}\right]. \tag{J.20b}$$

Once we allow equilibrium wages to vary in the counterfactual, we must take a stance on how much these wages change in reaction to the domestic demand shock. This change in wages will be determined by each firm's labor demand and labor supply functions.

To derive each firm's labor demand function, we assume that firms' production function is Cobb-Douglas in a fixed input (e.g. capital) and labor, which is treated as fully flexible input. Appendix A shows that, in this case, we can write the labor demand of a firm i as

$$L_{it} = \frac{1}{\varphi_{it}}\frac{1}{1+\lambda}(Q_{it})^{1+\lambda} = \frac{1}{\varphi_{it}}\frac{1}{1+\lambda}(\tau_{sdt}Q_{idt} + \tau_{sxt}Q_{ixt})^{1+\lambda}. \tag{J.21}$$

Furthermore, keeping the monopolistic competition assumption introduced in section 7.1, we can write the output price of firm i at period t in the domestic and foreign markets as

$$P_{idt} = \frac{\sigma}{\sigma-1}\frac{\omega_{it}\tau_{sdt}}{\varphi_{it}}(\tau_{sdt}Q_{idt} + \tau_{sxt}Q_{ixt})^{\lambda}, \tag{J.22}$$

$$P_{ixt} = \frac{\sigma}{\sigma-1}\frac{\omega_{it}\tau_{sxt}}{\varphi_{it}}(\tau_{sdt}Q_{idt} + \tau_{sxt}Q_{ixt})^{\lambda}. \tag{J.23}$$

Combining equations (J.21) and (J.22), we can write

$$\begin{aligned} L_{it} &= \frac{\sigma-1}{\sigma}\frac{P_{idt}}{\omega_{it}\tau_{sdt}}\frac{1}{1+\lambda}(\tau_{sdt}Q_{idt}+\tau_{sxt}Q_{ixt}) \\ &= \frac{\sigma-1}{\sigma}\frac{1}{\omega_{it}}\frac{1}{1+\lambda}(P_{idt}Q_{idt}+P_{idt}\frac{\tau_{sxt}}{\tau_{sdt}}Q_{ixt}) \\ &= \frac{\sigma-1}{\sigma}\frac{1}{\omega_{it}}\frac{1}{1+\lambda}(P_{idt}Q_{idt}+\frac{P_{idt}}{P_{ixt}}\frac{\tau_{sxt}}{\tau_{sdt}}P_{ixt}Q_{ixt}). \end{aligned}$$

Furthermore, given equations (J.22) and (J.23), it holds that $(P_{idt}/\tau_{sdt})/(P_{ixt}/\tau_{sxt}) = 1$ and, thus, we can further rewrite the labor demand equation as

$$L_{it} = \frac{\sigma-1}{\sigma}\frac{1}{\omega_{it}}\frac{1}{1+\lambda}(P_{idt}Q_{idt}+P_{ixt}Q_{ixt}) = \frac{\sigma-1}{\sigma}\frac{1}{\omega_{it}}\frac{1}{1+\lambda}R_{it}. \tag{J.24}$$

Given this labor-demand equation, we can write the log-difference between the boom and the bust periods in the labor demanded by a firm i as

$$\ln\left[\frac{L_{i1}}{L_{i0}}\right] + \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right] = \ln\left[\frac{R_{i1}}{R_{i0}}\right]. \tag{J.25}$$

We assume that every firm i faces an isoelastic labor supply curve. Denoting the inverse labor elasticity as ψ, we can thus write

$$\ln\left[\frac{L_{i1}}{L_{i0}}\right] = \frac{1}{\psi}\ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]. \tag{J.26}$$

This equation assumes that the shifter of the firm-specific labor supply curve does not vary over time. Our analysis is however robust to the presence of a shifter that changes over time, as long as it is invariant to the counterfactual changes in sector-specific aggregate demand whose impact we evaluate.

Combining the firm's labor demand and supply functions in equations (J.25) and (J.26), we obtain the following expression for the log change in firm-level wages as a function of the log change in firm-level total revenues:

$$(1+\frac{1}{\psi})\ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right] = \frac{1+\psi}{\psi}\ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right] = \ln\left[\frac{R_{i1}}{R_{i0}}\right],$$

or, equivalently,

$$\ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right] = \frac{\psi}{1+\psi}\ln\left[\frac{R_{i1}}{R_{i0}}\right].$$

In the case in which the sectoral domestic shocks in the bust period differ between actual and counterfactual scenarios, we can generally write the counterfactual change in firm-specific wages and revenues as

$$\ln\left[\frac{\omega'_{i1}}{\omega_{i0}}\right] = \frac{\psi}{1+\psi}\ln\left[\frac{R'_{i1}}{R_{i0}}\right]. \tag{J.27}$$

Plugging equation (J.27) into equations (J.20a) and (J.20b), we obtain

$$\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right]=\ln\left[\frac{B_{sx1}}{B_{sx0}}\right]+(\sigma-1)\ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right]+\frac{(\sigma-1)}{1+\lambda}\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right]-(\sigma-1)\ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right]$$
$$+\sigma\ln\left[\frac{P_{sx1}}{P_{sx0}}\right]-\frac{(\sigma-1)}{1+\lambda}\left(\lambda+\frac{\psi}{1+\psi}\right)\ln\left[\frac{R'_{i1}}{R_{i0}}\right], \quad \text{(J.28a)}$$
$$\ln\left[\frac{R'_{id1}}{R_{id0}}\right]=\ln\left[\frac{B'_{sd1}}{B_{sd0}}\right]+(\sigma-1)\ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right]+\frac{(\sigma-1)}{1+\lambda}\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right]-(\sigma-1)\ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right]$$
$$+\sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\right]-\frac{(\sigma-1)}{1+\lambda}\left(\lambda+\frac{\psi}{1+\psi}\right)\ln\left[\frac{R'_{i1}}{R_{i0}}\right]. \quad \text{(J.28b)}$$

Implementing the same steps as in the baseline counterfactual exercise (see Appendix J.1), we can further rewrite the expressions in equations (J.28a) and (J.28b) as

$$\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right]=\ln\left[\frac{R_{ix1}}{R_{ix0}}\right]$$
$$-\frac{(\sigma-1)}{1+\lambda}\left(\lambda+\frac{\psi}{1+\psi}\right)\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0}+\frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right)-\ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \quad \text{(J.29a)}$$
$$\ln\left[\frac{R'_{id1}}{R_{id0}}\right]=\ln\left[\frac{B'_{sd1}}{B_{sd0}}\left(\frac{B_{sd1}}{B_{sd0}}\right)^{-1}\right]+\sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right]+\ln\left[\frac{R_{id1}}{R_{id0}}\right]$$
$$-\frac{(\sigma-1)}{1+\lambda}\left(\lambda+\frac{\psi}{1+\psi}\right)\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0}+\frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right)-\ln\left[\frac{R_{i1}}{R_{i0}}\right]\right]. \quad \text{(J.29b)}$$

Our evaluation of the impact on firms' exports and domestic sales of counterfactual relative changes sector-specific demand shifters while allowing for firm-specific wages that change as firms move along their supply curves relies on three sets of equations: that defined by equation (J.11), and those defined by equations (J.29a) and (J.29b), which are respectively a generalization of the expressions in equations (J.4) and (J.5). The system of equations formed by equation (J.11) and equations (J.29a) and (J.29b) depends on the following parameters

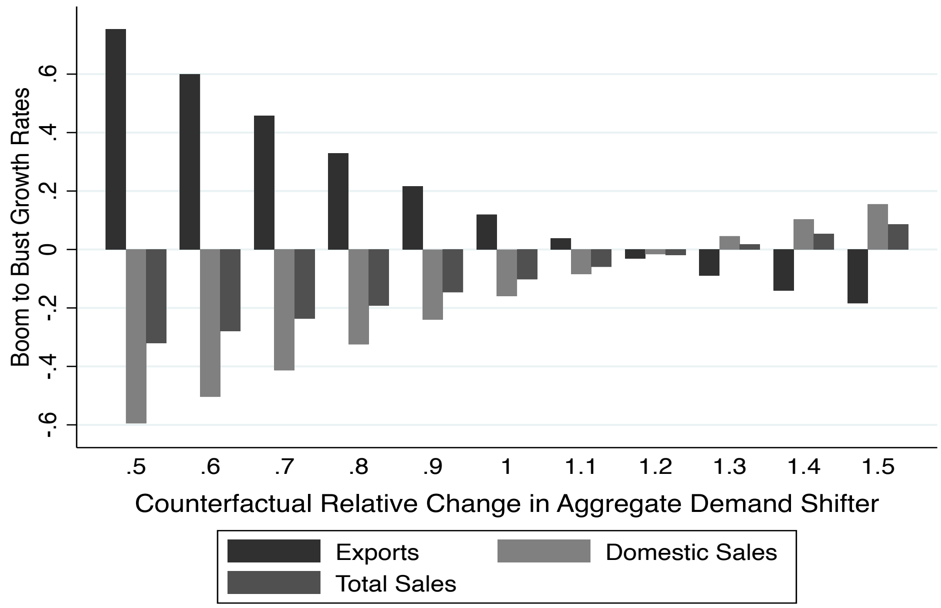
$$\left\{\sigma,\frac{(\sigma-1)}{1+\lambda},\lambda,\frac{\psi}{1+\psi}\right\}.$$

As in the baseline calibration, we set $\sigma=5$. Given this value of σ and the estimate $(\sigma-1)\lambda/(1+\lambda)=2.374$, which corresponds to the estimate reported in column 3 of Table 10, we can solve for λ obtaining an estimate $\lambda=1.46$. Concerning the value of $\psi/(1+\psi)$, we present results for three different values of this parameter: our baseline calibration (see section VII) asumes $\psi=0$; and we present here results for $\psi/(1+\psi)=0.5$ (i.e., $\psi=1$) and $\psi/(1+\psi)=1$ (i.e., $\psi=\infty$). The baseline calibration thus considers a case in which changes in firms' labor demand impact firms' labor usage, but not their wages; conversely, the calibration that sets $\psi/(1+\psi)=1$ considers a case in which changes in firms' labor demand impact firms' wages, but not their labor usage.

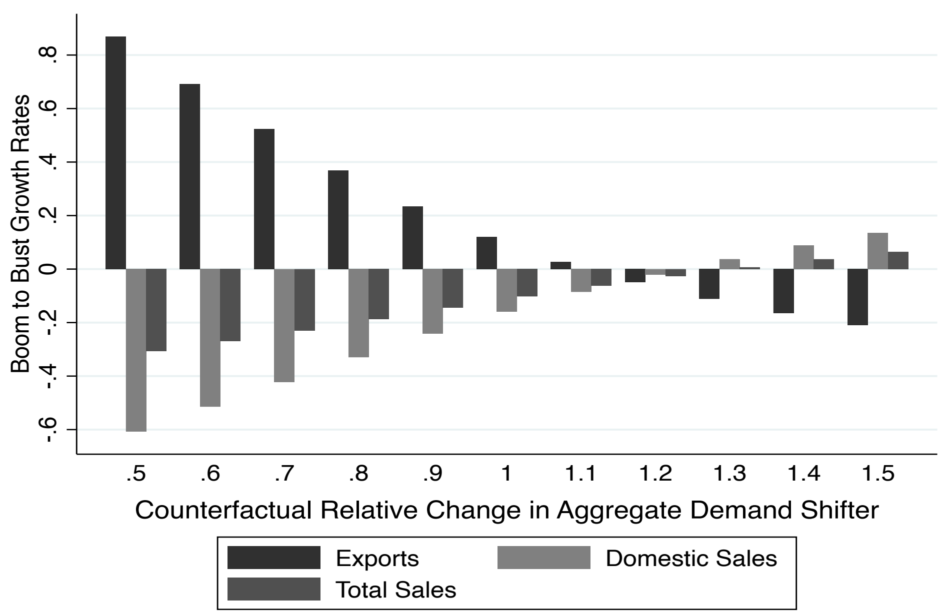
A comparison of the results in Figure 4 (see section VII in the main text) with those in Figure J.3 reveal that lower elasticities of labor supply imply that firms' exports react more to changes in domestic demand shifters. Thus, the lower the elasticity of the labor supply function that a firm faces, the more important the vent-for-surplus mechanism is. For example, if the demand shifter in the bust had been only 50% of its actual level (i.e., the boom-to-bust negative demand shock had been larger than observed in the data), our model predicts that aggregate exports would have grown in 60.1% if firm-specific wages do not adjust (see Figure 4), and in more than 80% if all

Figure J.3: Impact of Aggregate Demand Shocks For Different Labor Elasticities

(a) Labor Supply with Unit Elasticity ($\psi/(1+\psi)=0.5$)



(b) Completely Inelastic Labor Supply ($\psi/(1+\psi)=1$)



Notes: The horizontal axis indicates the value of Γ_B. The export and domestic sales growth rates indicated in the vertical axis correspond to those predicted by equations (J.11), (J.28a) and (J.28b). Given these counterfactual growth rates in export and domestic sales, we compute the counterfactual growth rate in total sales as $(R'_{ix1}/R_{ix0})\chi_{i0} + (R'_{id1}/R_{id0})(1-\chi_{i0})$.

the adjustment a firm's labor demand takes place through wages (and not through the quantity of labor that firms use, see Panel (b) in Figure J.3). Intuitively, the lower the elasticity of labor supply, the larger the drop in equilibrium wages in reaction to a drop in domestic demand and, thus, the larger the gains in competitiveness of Spanish firms in foreign markets (i.e., the larger the drop in the marginal cost of Spanish firms).

The assumptions we impose on the extent to which firms' wages react to firms' sales impact our quantification of the relevance of the vent-for-surplus mechanism. More specifically, they impact the value of Γ_B for which the corresponding model predicts a drop in total sales equal to $6.04\% = (1-41\%) \times 10.23\%$; i.e., the drop in total sales that we would have observed if aggregate demand shifters had not changed between the boom and the bust periods (see section VII in the main text for additional details). In our baseline calibration with invariant wages, this value of Γ_B was 1.09 (which implies that the actual drop in the aggregate demand shifters between boom and bust periods was $1-1/1.09 = 8.26\%$). In the model with unit-elastic labor supply functions (i.e., $\psi = 1$), this value of Γ_B is 1.1 (which implies a drop in the aggregate demand shifters of $1-1/1.1 = 9.09\%$). Finally, in the model with completely inelastic labor supply functions (i.e., $\psi = \infty$), this value of Γ_B is 1.11 (which implies a drop in the aggregate demand shifters of $1-1/1.1 = 9.91\%$).

For the corresponding calibrated values of Γ_B, our baseline results with fully elastic labor supply predict a growth in aggregate exports of 5.79%, those that assume a firm-specific labor supply with unit elasticity predict a growth in aggregate exports of 3.81%, and those that assume a perfectly inelastic labor supply predict a growth in aggregate exports of 1.88%. Consequently, as the observed growth in exports was 11.99%, our analysis indicates that the vent-for-surplus mechanism explains: $(11.99\% - 5.79\%)/11.99\% = 51.71\%$ of the total growth in exports when wages do not adjust; $(11.99\% - 3.81\%)/11.99\% = 68.22\%$ of the total growth in exports when the firm-specific labor supply has a unit elasticity; and $(11.99\% - 1.88\%)/11.99\% = 84.33\%$ of the total growth in exports when all the adjustment in firms' labor demand happens through wages.

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