Venting Out: Exports during a Domestic Slump*

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Abstract

We exploit plausibly exogenous geographical variation in the reduction in domestic demand caused by the Great Recession in Spain to document the existence of a robust, within-firm negative causal relationship between demand-driven changes in domestic sales and export flows. Spanish manufacturing firms whose domestic sales were reduced by more during the crisis observed a larger increase in their export flows, even after controlling for firms' supply determinants (such as labor costs). This negative relationship between demand-driven changes in domestic sales and changes in export flows illustrates the capacity of export markets to counteract the negative impact of local demand shocks. We rationalize our findings through a standard heterogeneous-firm model of exporting expanded to allow for non-constant marginal costs of production. Using a structurally estimated version of this model, we conclude that the firm-level responses to the slump in domestic demand in Spain could have accounted for around one-half of the spectacular increase in Spanish goods exports (the so-called 'Spanish export miracle') over the period 2009-13.

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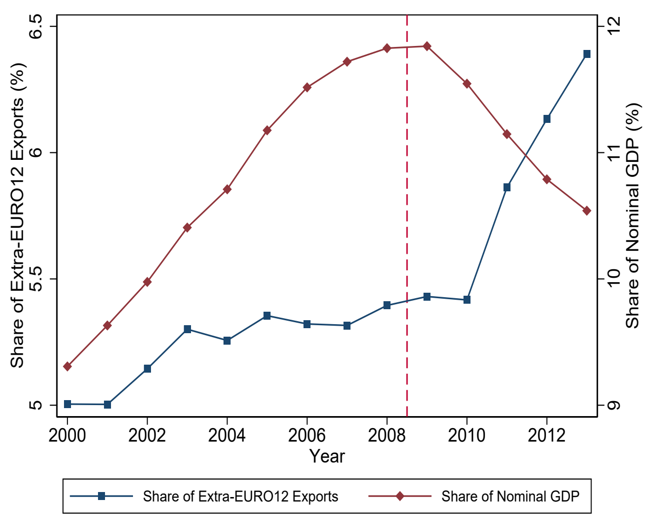
1 Introduction

The Great Recession of the late 2000s and early 2010s shook the core of many advanced economies. Few countries experienced the consequences of the global downturn as intensively as the Southern economies of the European Monetary Union (EMU) did. Spain is a case in point. From its peak in 2008, Spain's real GDP fell by an accumulated 8.9% in the following five years, until bottoming out in 2013. During the same period, private final consumption contracted by 14.0%. Portugal and Greece also experienced marked contractions between 2008 and 2013, with their GDPs shrinking by 7.9% and 26.3%, respectively.

Despite these severe domestic slumps, merchandise exports in these economies demonstrated a remarkable resilience and partly contributed to mitigating the effects of the Great Recession. In the Spanish case, after tumbling by 11.5% in real terms during the global trade collapse of 2008-2009, Spanish merchandise exports quickly recovered and grew by 30.7% in real terms between 2009 and 2013. Overall, real Spanish merchandise exports grew by an accumulated 15.6% during the 2008-2013 period, while real merchandise exports in the rest of the euro area increased by only 6.8% during the same years. As a result, and as shown in Figure 1, the share of euro area merchandise exports to non-euro area countries accounted for by Spain increased markedly during this period (especially in 2010-13), despite the contemporaneous decline in the relative weight of Spain's GDP in the euro area's GDP. Very similar patterns are observed for the cases of Portugal and Greece.[1]

Figure 1: The Spanish Export Miracle



At first glance, this remarkable export performance appears to be consistent with the goals of the type of "internal devaluation" processes advocated by international organizations (such as the IMF, the ECB or the European Commission) since the onset of the crisis. According to these

[1]In Appendix D.1, we replicate Figure 1 for Portugal, Greece, Ireland, Italy, Germany, France, and the Netherlands. For Portugal and Greece, and less clearly for Germany and France, we observe a negative relationship between their GDP shares in the euro area and their shares in euro area exports of goods to other countries. See Appendix D.1 for a description of the data sources underlying these figures.

organizations, wage moderation coupled with a set of structural reforms (most notably labor market reforms) were expected to lead to a fall in relative unit labor costs, allowing Southern European firms to reduce their relative export prices and increase their market shares abroad. Nevertheless, in the Spanish case, the adjustment in labor costs achieved via these policies is estimated to have been modest up to 2013, and this channel is believed to have had a limited contribution to export growth over the period 2010-13 (see, for instance, IMF, 2015, 2018; Salas, 2018).

What explains then the remarkable export growth in Spain over the period 2010-13? An often-invoked explanation relates the growth in exports *directly* to the collapse in domestic demand.[2] According to this hypothesis, the unexpected demand-driven reduction in firms' domestic sales, in combination with the irreversibility of certain investments in inputs, freed up capacity that these firms used to serve customers abroad. More precisely, this explanation posits that, as domestic demand fell, Spanish firms were able to cut their short-run marginal costs by reducing their usage of flexible inputs (e.g., temporary workers and materials) relative to their usage of fixed inputs (e.g., physical capital and permanent workers). This fall in short-run marginal costs translated into a gain in competitiveness in foreign markets and, consequently, to an increase in exports.[3]

This alternative explanation resonates with the "vent-for-surplus" theory of the benefits of international trade, which has a long tradition in economics dating back to Adam Smith.[4] Despite its intuitive nature and distinguished lineage, the link between a domestic slump and export growth is hard to reconcile with modern workhorse models of international trade. The reason for this is that these canonical models – including those emphasizing economies of scale – assume that firms face constant marginal costs of production, an assumption that implies that demand shocks in one market do not affect a firm's sales in another market.

In this paper, we leverage Spanish firm-level data from 2002 to 2013, to study the empirical relevance of the "vent-for-surplus" mechanism. To do so, and given the contrasting macroeconomic behavior in the pre- and post-crisis period, we first divide our sample into a "boom" period (2002-08) and a "bust" period (2009-13), and aim to measure the extent to which, at the firm level, a demand-driven decline in the domestic sales in the bust period relative to the boom period is associated with an increase in export sales over the two periods.

To guide us in the specification of an empirical strategy to estimate the causal impact of demand-driven changes in domestic sales on exports, we first rely on a commonly used model of firms' export behavior à la Melitz (2003). This framework helps us identify several empirical challenges that one

[2] See "La exportación como escape" in El País, 1/16/2016, for a journalistic account in Spanish with some specific case studies (https://elpais.com/economia/2016/01/14/actualidad/1452794395_894216.html). Further firm-level examples are provided in the more recent "El milagro exportador español" in El País, 5/27/2018 (http://elpais.com/economia/2018/05/25/actualidad/1527242520_600876.html).

[3] One can interpret this explanation as encompassing any mechanism that makes firms' short-run marginal cost curves increasing and that, thus, links the drop in firms' domestic demand to a downward movement *along* their supply curves. This effect is distinct from that of an "internal devaluation", which is associated with a downward *shift* in firms' marginal cost or supply curves caused by a reduction in the price of factors of production.

[4] In *The Wealth of Nations* (1776) Book II, Chapter V, Smith writes "When the produce of any particular branch of industry exceeds what the demand of the country requires, the surplus must be sent abroad, and exchanged for something for which there is a demand at home. Without such exportation, a part of the productive labour of the country must cease, and the value of its annual produce diminish." The term "vent-for-surplus" was introduced by John Stuart Mill in his *Principles of Political Economy* (1848) and popularized by Williams (1929) and Myint (1958).

encounters when measuring the relevance of the "vent-for-surplus" mechanism.[5] We draw two main conclusions from our theoretical analysis. First, as long as firms' marginal cost shifters (i.e., firms' productivity and production factor costs) are not perfectly observable – and their unobserved component is not fully captured by various fixed effects – there will tend to be a *positive* spurious correlation between domestic sales and exports that is not informative about the impact of demand-driven changes in the former on the latter. Second, an instrumental variable approach identifies the causal impact of demand-driven changes in domestic sales on exports as long as the instrument satisfies two conditions: (i) it is a good predictor of the domestic sales of Spanish firms, and (ii) it is not correlated with firms' unobserved marginal cost or export-demand shifters.

With these considerations in mind, we first show that, at the firm level, a decline in the domestic sales in the bust period relative to the boom period is associated with an increase in export sales over the two periods. When measuring this association, we control for "boom-to-bust" changes in observed marginal cost shifters (i.e., measures of factor prices and productivity) to account for potential internal devaluation effects. However, it is likely that there are additional marginal cost shifters that we do not observe and, thus, the estimated negative association between boom-to-bust changes in domestic sales and exports likely underestimates the true positive impact on the firm's exports of demand-driven reductions in its domestic sales.

To better estimate the export impact of demand-driven changes in domestic sales, we rely on two instrumental variables that exploit the fact that the Great Recession affected different geographical areas in Spain differentially. Both instrumental variables rely on municipality-level registration data on a major household durable consumption item, vehicles, as a proxy for the extent to which the Great Recession affected the demand for Spanish manufacturing goods in each municipality. The use of vehicle purchases as a proxy for 'local demand' is justified by an extensive literature in empirical macroeconomics documenting that consumption of durable goods such as vehicles is strongly procyclical (see, for instance, the survey by Stock and Watson, 1999). More recently, Mian et al. (2013), Hausman et al. (2019), and Waugh (2019) have also documented a strong link between wealth (or income) shocks and vehicle consumption. While both of our instrumental variables rely on vehicle registration as a proxy for local demand, they differ in the assumed exposure of each firm to changes in local demand across different Spanish municipalities.

To measure the extent to which firms were differentially exposed to the demand shocks that different Spanish municipalities experienced in the years around the Great Recession, we use firm-to-municipality manufacturing sales data (from tax records within Spain) for the year 2006. A clear feature of this data is that shipments by Spanish manufacturing firms are extremely localized, consistently with the facts documented in Hillberry and Hummels, (2008) for the U.S. and Diaz-Lanchas et al. (2019) for Spain. Firms' shipments within their municipality of location are on average five times larger than those to any other municipality, after controlling for bilateral distance and population. Consequently, our first instrumental variable uses boom-to-bust changes in the

[5] The Melitz (2003) model assumes that firms face constant marginal costs of production, implying the null hypothesis of a zero effect of demand-driven changes in domestic sales on exports. However, as we show in section 7, the lessons we learn from this model in terms of the econometric challenges one faces when evaluating the "vent-for-surplus" mechanism are also applicable to more general models that feature increasing marginal costs of production.

stock of vehicles per capita in a municipality as instrument for the boom-to-bust changes in the Spain-wide sales of firms located in that municipality.

A second clear pattern in the firm-to-municipality manufacturing sales data is that a large fraction of Spanish firms ship outside of their municipality the bulk of their domestic sales. Consequently, we also use a second, model-based, instrumental variable that, for each firm, equals a weighted sum of our proxy for the municipality-specific local demand shocks, using as weights gravity-based estimates of the forces the determine the size of the bilateral trade flows between any two Spanish municipalities. We compute these gravity-based estimates using municipality-to-municipality sales data (aggregated from the firm-to-municipality tax records). These estimates reveal a significant amount of "home bias" *within Spain*, with shipments declining in distance with an elasticity of around -0.4 even after controlling for the discontinuity in sales observed when shipping outside a firm's municipality. To ensure that this alternative instrument exploits distinct variation than the first one, we assign a zero weight to the municipality where the firm is located: this second instrument is thus a weighted-sum of the stock of vehicles per capita in all municipalities *other than* the municipality in which the firm is located.

The first-stage estimates for both of our instrumental variables suggest that both of them are strong (with F-statistics above 30 in both cases). Armed with these first-stage results, we then show that a larger *demand-driven* drop in domestic sales in the bust period relative to the boom period is associated with a significantly larger growth in export sales from boom to bust (conditional on exporting in both periods). Furthermore, the IV estimates based on our two instruments are significantly larger than the OLS ones. This is consistent with the biases predicted by our baseline Melitz-style model in the plausible scenario in which our covariates only imperfectly control for a firm's supply determinants. Specifically, depending on which of the two instrumental variables described above we use, we obtain estimates of the intensive-margin elasticity of exports to domestic sales in the neighborhood of -1.3 or -1.6, while the OLS one is around -0.3. Besides our baseline two-period specification, which exploits changes in exports, domestic sales and our instrumental variables between the boom and bust periods, we also present panel estimates that exploit the annual frequency of our data, and which allow for the inclusion of municipality-specific time trends as additional controls. Although our annual panel OLS estimates are very similar to those obtained in the boom-to-bust specification, our IV estimates are not, the reason being that our instruments lose their predictive power at the annual frequency. Nevertheless, when considering rolling averages over two or three years, our instruments regain strength and the resulting IV estimates resemble those in the boom-to-bust specifications even when municipality-specific trends are accounted for.

A potential challenge to our identification approach is that the boom-to-bust changes in our instruments may be correlated with the extent to which unobserved shifters of the firm's marginal cost curve changed in the bust period relative to the boom period. This is arguably less problematic for our second instrument, which does not use information on demand changes in the firm's own municipality when constructing the instrument, but even in that case one may still be concerned about spatial correlation in supply shocks posing a threat to identification. With that in mind, we provide in section 6 additional pieces of evidence that address some specific sources of endogeneity that could affect the validity of our instruments.

First, an identification threat arises if differences in the severity of the contraction in vehicle purchases across Spanish municipalities are not exclusively a reflection of differences in demand shocks, but also partly a reflection of unobserved production costs affecting car manufacturers. Specifically, if a significant share of vehicles is sold in the vicinity of where they are produced, municipalities that concentrate a significant share of firms operating in the auto industry could observe a correlation in the boom-to-bust changes in production costs and nearby purchases of new vehicles. Our results are robust to this identification threat. Both the relevance of our instrument and the finding of a sizable negative elasticity between domestic sales and exports are robust to excluding firms that could have been impacted by production costs shocks affecting car manufacturers.

Second, although we control for firm-specific average wages in all specifications, compositional changes in firms' workforce may have caused changes in effective labor costs that our wage measure does not capture correctly. Indeed, an important feature of the Spanish labor market is the division of the workforce into permanent and temporary workers, the latter group being typically less productive than the former (see Dolado et al., 2002). We find, however, that the elasticity of exports with respect to domestic sales remains largely unaffected when we control for the firm's change in the share of temporary workers. Similarly, controlling for the change in financial costs experienced by exporters or for proxies of trade credit available to firms does not materially affect our main estimates. In addition, we also explore in section 6 the robustness of our results to alternative constructions of our instrumental variable that rely on different proxies for the exposure of a firm to demand shocks in municipalities other than its municipality of location, as well as to alternative proxies for a firm's total factor productivity.

Having established a causal link between changes in domestic demand and exports, we generalize our baseline model à la Melitz (2003) to allow for non-constant marginal costs of production. We rationalize this cost structure by including a pre-determined and fixed factor into the firm's production function, and show that the curvature of the firm's marginal cost function is related to the elasticity of output with respect to all flexible factors. Furthermore, we demonstrate how to estimate the curvature of the marginal cost function using a simple variant of our IV estimator. Consistently with our micro-foundation, we find that our estimate of this curvature is smaller for firms whose output elasticity with respect to flexible factors is larger, though the statistical significance of this result is sensitive to which factors one classifies as "flexible".

Finally, we employ our model with increasing marginal costs and the corresponding IV estimates to quantitatively evaluate the importance of the "vent-for-surplus" mechanism in explaining the 2009-13 export miracle in Spain. More specifically, we implement a variance-decomposition exercise to determine the extent to which the domestic slump in Spain was driven by demand versus supply shocks. We then use our model to predict the boom-to-bust growth in Spanish exports that we would have observed if there had been no change in demand between the boom and bust periods. We find that, in this case, the growth in Spanish exports would have been 51.71% smaller than what we observe in the data and, thus, we conclude that slightly more than half of the Spanish export miracle of the period 2009-2013 can be attributed to the "vent-for-surplus" mechanism.

Our paper connects with several branches of the literature. As mentioned above, we relate the Spanish export miracle to Adam Smith's "vent-for-surplus" theory. The international trade literature has largely ignored this hypothesis as exemplified by the fact that we have only found one mention (in Fisher and Kakkar, 2004) of the term "vent-for-surplus" in all issues of the *Journal of International Economics*.[6] Nevertheless, there has been an active recent international trade literature focused on relaxing the assumption of constant marginal costs in the canonical (Melitz) model of firm-level trade, and has shown that, in the presence of increasing marginal costs, there is a natural substitutability between domestic sales and exports for which there is supporting empirical evidence. This literature includes the work of Vannoorenberghe (2012), Blum, Claro and Horstman (2013), Soderbery (2014), and Ahn and McQuoid (2017). Relative to this prior literature, our paper exploits plausibly exogenous variation in demand during a particularly salient episode to identify the causal effect of a demand-driven drop in domestic sales on exports. Additionally, it provides an approach to identify and structurally estimate the slope of firms' short-run marginal cost curves. Relatedly, in contemporaneous work, Fan et al. (2020) exploit variation in the extent to which Chinese authorities enforce the collection of value-added taxes to establish a negative causal link between the profitability of domestic sales and firm-level exports. Conversely, using French data over the period 1995-2001, Berman et al. (2015) document a *positive* causal effect of changes in firm-level exports on firm-level domestic sales. Their identification strategy (based on exogenous variation in foreign demand conditions) is quite distinct from ours and so is their setting, since 1995-2001 was a tranquil period of sustained economic growth in France. In Appendix H.1, we use data on Spanish firms for the period 2002-07 to perform an analysis analogous to that in Berman et al. (2015), and we find no evidence supporting the positive causal relationship between exports and domestic sales that these authors previously found; on the contrary, for most specifications, we find a *negative* causal effect of (plausibly) exogenous changes in exports on domestic sales, in line with our core finding of substitution between exports and domestic sales.[7]

The rest of the paper is structured as follows. In section 2, we lay out a baseline model of firm behavior in the spirit of Melitz (2003). In section 3, we introduce our firm-level data and, in section 4, we develop our core instrumental variable estimation approach. Our main results are presented in section 5, while we present additional evidence in favor of the "vent-for-surplus" mechanism in section 6. In section 7, we generalize our baseline model to allow for non-constant marginal costs. In section 8, we use this extended framework to quantify the contribution of the "vent-for-surplus" channel in the growth of Spanish exports. We offer some concluding remarks in section 9.

2 Benchmark Model: Estimation Guidelines

As indicated in the Introduction, we aim to estimate the causal impact of within-firm demand-driven changes in domestic sales on firm-level exports. To guide our empirical analysis, we first consider the implications of a model of exporting with heterogeneous firms along the lines of Melitz

[6] A broader search to include top general-interest journals identified Neary and Schweinberger (1986).

[7] Our paper also relates to a prior literature describing the behavior of firm-level exports in Spain around the Great Recession, including Antràs (2011), Myro (2015), Eppinger et al. (2017), and De Lucio et al. (2017a, 2017b).

(2003), the canonical model of firm-level exports in the recent international trade literature. This model features the standard assumption of constant marginal costs. After presenting evidence contradictory with this assumption, we will develop in section 7 an extension of this benchmark model that allows for non-constant marginal costs. Crucially, the lessons we learn in this section about the properties of different estimators will also apply in the more general model.

2.1 Benchmark Model: Estimating Equation

We index manufacturing firms producing in Spain by i, the sectors to which firms belong by s, and the two potential markets in which they may sell by $j = \{d, x\}$, with d denoting the domestic market and x denoting the export market. In principle, both the domestic and export market are an aggregate of several destinations, but due to data limitations, we focus in the main text on this dichotomous case (we develop a multi-destination extension of our model in Appendix E.2).

In any given period, firm i faces the following isoelastic demand in market j,

$$Q_{ij} = \frac{P_{ij}^{-\sigma}}{P_{sj}^{1-\sigma}} E_{sj} \xi_{ij}^{\sigma-1}, \qquad \sigma > 1, \tag{1}$$

where Q_{ij} denotes the number of units of output of firm i demanded in market j if it sets a price P_{ij}, P_{sj} is the sector s price index in j, E_{sj} is the total sectoral expenditure in market j, and ξ_{ij} is a firm-market specific demand shifter.

Firm i's total variable cost of producing Q_{ij} units of output for market j is given by

$$c_{ij} Q_{ij} \qquad \text{with} \qquad c_{ij} \equiv \tau_{sj} \frac{1}{\varphi_i} \omega_i, \tag{2}$$

where c_{ij} denotes the marginal cost to firm i of selling one unit of output in market j, τ_{sj} denotes an iceberg trade cost, φ_i denotes firm i's productivity, and ω_i is the firm-specific cost of a bundle of inputs. Additionally, we assume that firm i needs to pay an exogenous fixed cost F_{ij} to sell a positive amount in market j.

Firm i chooses optimally the quantity offered in each market j, Q_{ij}, taking the price index, P_{sj}, and the size of the market, E_{sj}, as given. As the marginal production cost is independent of the firm's total output and the per-market fixed costs are independent of the firm's participation in other markets, the optimization problem of the firm is separable across markets. Specifically, conditional on selling to a market j, firm i solves the following optimization problem

$$\max_{Q_{ij}} \left\{ Q_{ij}^{\frac{\sigma-1}{\sigma}} P_{sj}^{\frac{1-\sigma}{\sigma}} E_{sj}^{\frac{1}{\sigma}} \xi_{ij}^{\frac{\sigma-1}{\sigma}} - \tau_{sj} \frac{1}{\varphi_i} \omega_i Q_{ij} \right\},$$

and sales by firm i to market j are thus $R_{ij} = P_{ij} Q_{ij} = \kappa \left((\xi_{ij} \varphi_i)/(\tau_{sj} \omega_i) \right)^{\sigma-1} E_{sj} P_{sj}^{\sigma-1}$, where κ is a function of σ. For the case of exports ($j = x$), and taking logs, we can rewrite this expression as:

$$\ln R_{ix} = \ln \kappa + (\sigma - 1)\left(\ln \xi_{ix} + \ln \varphi_i - \ln \omega_i\right) - (\sigma - 1)\left(\ln \tau_{sx} - \ln P_{sx}\right) + \ln E_{sx}. \tag{3}$$

The bulk of our empirical analysis will compare firm-level export behavior in a *bust* period, relative to a *boom* period. With that in mind, and letting $\Delta \ln X$ denote the log change in the cross-year average value of X from boom to bust, we can express the log change in exports from boom to bust as

$$\Delta \ln R_{ix} = (\sigma - 1) \left[\Delta \ln \xi_{ix} + \Delta \ln \varphi_i - \Delta \ln \omega_i \right] - (\sigma - 1) \left(\Delta \ln \tau_{sx} - \Delta \ln P_{sx} \right) + \Delta \ln E_{sx}. \tag{4}$$

To transition to an estimating equation, we model the change in firm-specific foreign demand, productivity and input bundle cost as follows:

$$\begin{aligned} \Delta \ln \xi_{ix} &= \xi_{sx} + u^{\xi}_{ix}, \\ \Delta \ln \varphi_i &= \varphi_s + \delta_{\varphi} \Delta \ln \varphi^*_i + u^{\varphi}_i, \\ \Delta \ln \omega_i &= \omega_s + \delta_{\omega} \Delta \ln \omega^*_i + u^{\omega}_i. \end{aligned} \tag{5}$$

Note that we are decomposing these terms into (i) a sector fixed effect, (ii) an *observable* part of these terms for the case of productivity (φ^*_i) and for input bundle cost (ω^*_i), and (iii) a residual term. We can thus re-write equation (4) as:

$$\Delta \ln R_{ix} = \gamma_{sx} + (\sigma - 1) \, \delta_{\varphi} \Delta \ln \varphi^*_i - (\sigma - 1) \, \delta_{\omega} \Delta \ln \omega^*_i + \varepsilon_{ix}, \tag{6}$$

where $\gamma_{sx} \equiv (\sigma - 1) \left[\xi_{sx} + \varphi_s - \omega_s - \ln \tau_{sx} + \ln P_{sx} \right] + \ln E_{sx}$, and where

$$\varepsilon_{ix} = (\sigma - 1) \left[u^{\xi}_{ix} + u^{\varphi}_i - u^{\omega}_i \right]. \tag{7}$$

Following analogous steps as above, we derive an expression for the change in domestic sales:

$$\Delta \ln R_{id} = \gamma_{sd} + (\sigma - 1) \, \delta_{\varphi} \Delta \ln \varphi^*_i - (\sigma - 1) \, \delta_{\omega} \Delta \ln \omega^*_i + \varepsilon_{id}, \tag{8}$$

where $\gamma_{sd} \equiv (\sigma - 1) \left[\xi_{sd} + \varphi_s - \omega_s - \ln \tau_{sd} + \ln P_{sd} \right] + \ln E_{sd}$, and where

$$\varepsilon_{id} = (\sigma - 1) \left[u^{\xi}_{id} + u^{\varphi}_i - u^{\omega}_i \right]. \tag{9}$$

We use equations (6) through (9) to generate predictions for the asymptotic properties of several estimators of the response of log exports to demand-driven changes in log domestic sales. The assumption of constant marginal costs implies that, according to this baseline model, the parameter of interest is zero: demand-driven changes in $\ln R_{id}$ have no causal effect on $\ln R_{ix}$. However, many estimators of the impact of log domestic sales on log exports based on observational data will differ from zero and, thus, be biased, even in large samples. We discuss here the asymptotic properties of different OLS and IV estimators.

Consider first using OLS to estimate the parameters of the following regression, which includes the change in log domestic sales as an additional covariate in equation (6):

$$\Delta \ln R_{ix} = \gamma_{sx} + (\sigma - 1)\,\delta_{\varphi} \Delta \ln \varphi_i^* - (\sigma - 1)\,\delta_{\omega} \Delta \ln \omega_i^* + \beta \Delta \ln R_{id} + \varepsilon_{ix}. \tag{10}$$

From equations (7), (9), and (10), the probability limit of the OLS estimator of the coefficient on domestic sales can be written as

$$plim(\hat{\beta}_{OLS}) = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}_{id})}{var(\Delta \ln \mathcal{R}_{id})} = \frac{cov(u_{ix}^{\xi} + u_i^{\varphi} - u_i^{\omega}, u_{id}^{\xi} + u_i^{\varphi} - u_i^{\omega})}{var(u_{id}^{\xi} + u_i^{\varphi} - u_i^{\omega})}, \tag{11}$$

where 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 and the observable covariates $\Delta \ln \varphi_i^*$, and $\Delta \ln \omega_i^*$.

We draw two conclusions from equation (11). First, if changes in productivity and production factor costs are not perfectly observable – and their unobserved component is not fully captured by the sector fixed effects – there will be a positive correlation between changes in exports and changes in domestic sales. Intuitively, unobserved productivity or factor cost changes will affect sales in the same direction in all markets in which a firm sells. In large samples, this will lead $\hat{\beta}_{OLS}$ to be positive and, thus, to be an upwards biased estimator of the impact of demand-driven changes in domestic sales on exports. Second, even if one proxies for changes in productivity and factor costs perfectly (i.e., $u_i^{\varphi} = u_i^{\omega} = 0$), in the presence of a non-zero correlation in the change in residual demand faced by firms in domestic and foreign markets (i.e., $\text{cov}(u_{ix}^{\xi}, u_{id}^{\xi}) \neq 0$), the estimator $\hat{\beta}_{OLS}$ will converge to a non-zero value. As this residual demand does not capture sector- and market-specific aggregate shocks (which are controlled by the sector fixed effects), it seems plausible that u_{ix}^{ξ} and u_{id}^{ξ} will be positively correlated, leading $\hat{\beta}_{OLS}$ again to be biased upwards.

Consider next using an IV estimator of the parameter β in equation (10). Specifically, consider instrumenting $\Delta \ln R_{id}$ with an observed covariate Z_{id} such that Z_{id} is either a proxy for $\Delta \ln \xi_{id}$ or has a causal impact on this firm-specific domestic demand shifter. In this case, the probability limit of the IV estimator of β is

$$plim(\hat{\beta}_{IV}) = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \mathcal{Z}_{id})}{cov(\Delta \ln \mathcal{R}_{id}, \mathcal{Z}_{id})} = \frac{cov(u_{ix}^{\xi} + u_i^{\varphi} - u_i^{\omega}, \mathcal{Z}_{id})}{cov(u_{id}^{\xi} + u_i^{\varphi} - u_i^{\omega}, \mathcal{Z}_{id})}, \tag{12}$$

where, as above, we use $\mathcal{Z}_{id}$ to denote the residual from projecting Z_{id} on a vector of sector fixed effects and the observable covariates $\Delta \ln \varphi_i^*$, and $\Delta \ln \omega_i^*$. The constant-marginal-cost model predicts that $\hat{\beta}_{IV}$ converges in probability to the true zero causal effect of demand-driven changes in domestic sales on exports as long as the variable $\mathcal{Z}_{id}$ satisfies two conditions: (a) it is correlated with the change in domestic sales of firm i after partialing out sector fixed effects as well as observable determinants of the firm's marginal cost; and (b) it is mean independent of the change in firm-specific unobserved productivity, u_i^{φ}, factor costs, u_i^{ω}, and export demand u_{ix}^{ξ}. As illustrated by the second equality in equation (12), an instrument can only verify conditions (a) and (b) if its effect on domestic sales works *exclusively* through the component of the change in domestic demand that is not accounted for by the sector fixed effects and the observable covariates included in the estimating equation, i.e., if it works exclusively through u_{id}^{ξ}.

Although our discussion above has centered around the role of unobserved supply and export demand factors in biasing estimates of β, Berman et al. (2015) emphasize that measurement error in both domestic sales and exports constitutes an additional source of possible bias when estimating the effect of exports on domestic sales (or vice versa). Because in many empirical settings – ours included – domestic sales are computed by subtracting exports from the firm's total sales, measurement error in firm total sales and exports will lead to a bias in the OLS estimate $\hat{\beta}_{OLS}$ that is likely to be of the opposite sign to that generated by the unobserved supply and export demand shocks accounted for by the residuals defined in equations (7) and (9). Consequently, as we detail in Appendix E.1 (see also Berman et al., 2015), negative values of $\hat{\beta}_{OLS}$ in large samples may be compatible with firms having constant marginal costs as long as the researcher's measures of either total sales or exports are affected by measurement error. Nevertheless, as we also show in Appendix E.1, if an instrument satisfies the same conditions (a) and (b) outlined above and is mean independent of the measurement error in exports, the IV estimator in equation (12) will still converge to zero in the presence of measurement error in total sales and exports.[8]

We have focused our discussion on the intensive margin of exports, namely the impact of domestic demand shocks on the level of exports conditional on exporting. In Appendix E.3, we show that an analysis of the extensive margin of exports modeled as a linear probability model delivers very similar insights. More specifically, when estimating the effect of demand-driven changes in domestic sales on the probability of exporting, even if the true effect were to be zero, one is likely to obtain a positive OLS estimate whenever productivity and production factor costs are not perfectly captured by sector fixed effects and observable controls, or whenever unobserved firm-specific demand shocks are positively correlated across domestic and export markets. An instrument satisfying conditions (a) and (b) above will continue to effectively remove these biases as long as it satisfies the additional condition of being mean independent of the part of the change in the firm's fixed cost of exporting not captured by sector fixed effects and marginal cost proxies.

3 Setting and Data

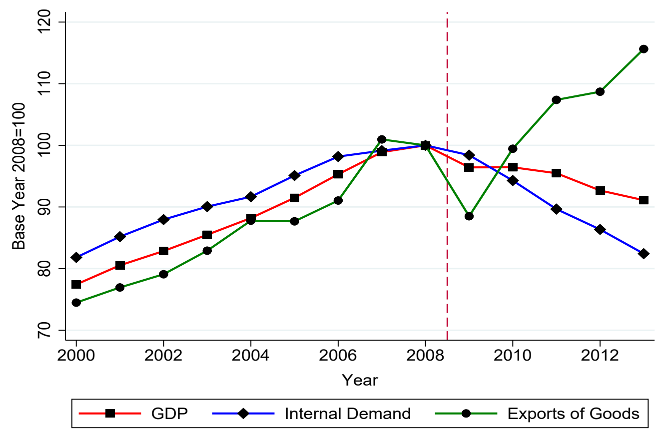
To construct a plausibly exogenous measure of the changes in domestic demand faced by firms, we exploit geographical variation within Spain in the severity of the Great Recession of the late 2000s and early 2010s. In this section, we describe the setting and data, and we defer a more detailed account of our identification strategy to section 4.

3.1 The Great Recession in Spain: Description

The macroeconomic history of Spain during the period 2000-2013 is a tale of a boom followed by a bust. As shown in Figure 2, between the year 2000 and the peak of the cycle in 2008, Spain's

[8]As mentioned above, we generalize our model in Appendix E.2 to incorporate multiple domestic and foreign markets. Theoretically, firms' choice over multiple export destinations may render the instruments $\mathcal{Z}_{id}$ invalid, even if they satisfy the conditions outlined in the main text. However, model simulations presented in Appendix E.2.3 illustrate that the resulting potential bias in the IV estimates is small for most parameter values.

Figure 2: The Great Recession in Spain



GDP and domestic demand grew by approximately 20% in real terms.[9] In the five subsequent years until 2013, domestic demand decreased to the level of the year 2000, while real GDP fell by an accumulated 8.9%.

The particularly severe impact of the Great Recession in Spain is largely explained by the fact that the economic boom of the early 2000s was primarily fueled by a real estate bubble. The construction sector accumulated an increasing share of GDP and employment.[10] For instance, in 2006, 658,000 new houses were built in Spain, a number corresponding to 80% of those built in Germany, Italy and the UK *combined* (EU Buildings Database). This real estate boom was in turn fostered by the increased availability of cheap credit to households, firms and real estate developers, which resulted from capital inflows related to the adoption of the euro in 2002 and the global savings glut (Santos, 2014). As a result, the ratio of mortgage credit to GDP went up from 40% in 2000 to 100% in 2008 (Basco et al. 2020). Importantly, the very high loan-to-value (LTV) ratios associated with residential mortgage credit were partly used by households to finance private consumption, particularly vehicle purchases (Masier and Villanueva, 2011).

The unraveling of the subprime mortgage market in the U.S. in the summer of 2007 had an immediate effect on the supply of credit in Spain. However, the effects were fully transmitted to the real economy only about one year later, coinciding with the fall of Lehman Brothers in September 2008 and the sudden stop in capital inflows (Basco et al., 2020). The recession officially started in the fourth quarter of 2008, and intensified during 2009 with a 3.6% annual drop in real GDP. The growth in the stock of vehicles in Spain, which had been stable at an average rate of 3.6% a year during the boom, suddenly came to a halt in 2008 (see Figure C.1 in Appendix C). In fact, in 2013, the national stock of vehicles in Spain was *lower* than in 2008 by around 52,000 units.

Importantly for the identification strategy we describe in the next section, the real estate boom and subsequent bust featured significant geographic variation, affecting mainly some parts of the

[9]Domestic demand is defined as final consumption expenditure by households and non-profit institutions serving households (NPISHs) plus investment plus acquisitions of public administrations minus imports.

[10]The share of total employment in construction peaked at 13.5% in 2007 and then collapsed to 5.4% by 2014, with a similar pattern for the contribution of this sector to Spain's GDP (12.4% in 2007 and 6.8% in 2014).

Mediterranean coast and medium-sized and large cities. As we shall document in section 4, this in turn translated into substantial geographic variation in the extent to which the Great Recession affected the domestic sales of Spanish firms.

3.2 The Spanish Export Miracle

As Figure 2 illustrates, the evolution of Spain's aggregate merchandise exports during the period 2008-13 was significantly different from that of aggregate domestic demand. After a significant 11.5% drop in real terms during the global trade collapse of 2008-09, aggregate exports grew during the period 2009-13 at an even faster rate than during the boom years. Specifically, while exports had grown by an accumulated 34% in the eight-year period 2000-08, they grew by a very similar 31% in just the four years between 2009 and 2013. This acceleration in export growth occurred at a time during which all indicators of domestic economic activity were showing a significant decline. As a consequence, the fall in real GDP was significantly smaller than the fall in domestic demand, and the ratio of exports of goods to GDP grew from 15.1% in 2009 to 23.3% in 2013. In Appendix D.2, we use the firms in our sample to describe the dynamics of the exports-to-sales ratio by sector.

One might wonder whether a depreciation in the euro could explain the growth in Spanish exports during the period 2009-13. Figure 1 in the Introduction shows that this could not have been the main explanation, as Spanish exports to non-euro area countries clearly outperformed those of other countries in the euro area (while Spain's GDP dropped faster than the euro area average).[11] It has also been argued that Spain underwent an internal devaluation during this period (through wage moderation starting in 2009, and via a labor market reform in 2012), but there is little evidence that these policies had a significant effect on relative production costs before 2012. For instance, unit labor costs in Spain were only 2.2% lower in 2012 than in their peak in 2009 (OECD Statistics). However, as we document in Appendix D.3, Spanish manufacturing export prices (unit values from product-level export data) fell relative to export prices in other euro area countries from the onset of the crisis, before Spanish unit labor costs had started to fall.

Motivated by these facts, we will hereafter focus on an exploration of the "vent-for-surplus" hypothesis, according to which the domestic slump, by freeing up production capacity, *directly* incentivized Spanish producers to sell their goods in foreign markets. More precisely, we hypothesize that the domestic slump led firms to move down along their short-run marginal-cost schedule, thereby lowering their export prices and leading them to gain market share in export markets.[12]

In principle, the 2009-13 growth in exports could have materialized along the intensive margin (with continuing exporters increasing their exports) or along the extensive margin (via net entry into the export market). Descriptive evidence suggests that the bulk of the growth was driven by the intensive margin. Using detailed Spanish Customs data, De Lucio et al. (2017a) find that net firm entry (i.e., new exporters net of firms quitting exporting) contributed a mere 14% to the

[11] Figure 1 also shows that most of the *relative* take-off of Spain occurred after 2010. The same is not true when looking at Spain's share in overall goods exports (including exports to euro area countries); in that case, Spain's share increased markedly already in 2009. This suggests that the increase in Spanish exports (relative to euro area countries) immediately following the Great Recession was largely driven by increased exports within the euro area.

[12] This is in contrast with the type of downward shift in marginal costs associated with internal devaluations.

export growth between 2008 and 2013, while the remaining 86% was driven by continuing exporters. Similarly, in our sample, we find that continuers contributed 91% of the growth in exports between the boom and the bust periods, and the extensive margin only accounted for 9% of export growth.[13]

3.3 Data Sources

Our data cover the period 2000-13 and come from various confidential administrative data sources. The first is the Commercial Registry (*Registro Mercantil Central*). It contains annual financial statements for around 85% of registered firms in the non-financial market economy in Spain. Among other variables, it includes information: sector of activity (4-digit NACE Rev. 2 code), 5-digit zip code of location, net operating revenue, material expenditures (cost of all raw materials and services purchased by the firm in the production process), labor expenditures (total wage bill, including social security contributions), number of employees (full-time equivalent), and total fixed assets. We provide more details regarding this dataset in Appendix B (see also Almunia et al., 2018).

The second dataset is the foreign transactions registry collected by the Bank of Spain (*Banco de España*). For both exports and imports, it contains transaction-level information on the fiscal identifier of the Spanish firm involved in the transaction, the amount transacted, the product code (SITC Rev. 4), the country of the foreign client, and the exact date of the operation (no matter when the payment was performed). Starting in 2008, however, the dataset's information on the product code and on the destination country became unreliable. The reason for this is that, to save on administrative costs, the entities reporting to the Bank of Spain were given the option of bundling a set of transactions together. In those cases, each entry reflects *only* the country of destination and product code of the largest transaction in that bundle (see Appendix B for more details). This feature of the dataset precludes us from studying exports at the firm-product-destination-year level during the crisis, but we can still reliably aggregate this transaction-level data to obtain information on total export volume by firm and year.

This international trade database has an administrative nature because *Banco de España* legally required financial institutions and external (large) operators to report this information for foreign transactions above a fixed monetary threshold. Until 2007, the minimum reporting threshold was fixed at 12,500 euros per transaction. Since 2008 until the end of the mandatory registry in 2013, information had to be reported for all transactions performed by a firm during a natural year as long as at least one of these transactions exceeded 50,000 euros. In order to homogenize the sample, for the period 2000 to 2007, we only record a positive export flow in a given year for firms that had at least one transaction exceeding 50,000 euros in that year (see Appendix B for more details). The foreign transactions registry collected by the Bank of Spain was discontinued in early 2014, which precludes us from extending our analysis past the year 2013.

In both datasets, a firm is defined as a business constituted in the form of a Corporation

[13]De Lucio et al. (2017a) also show that a third of the contribution of continuing exporters is due to entry into new destination countries and products, while the other two thirds are due to growth in existing product-country combinations. Unfortunately, the nature of the export data available to us does not allow us to explore the firm-level extensive margin at the product or destination country level.

(*Sociedad Anónima*), a Limited Liability Company (*Sociedad Limitada*), or a Cooperative (*Cooperativa*). We merge both datasets using each firm's fiscal identifier. Using the merged database, we define each firm's domestic sales as the difference between its total sales and its total exports.

To check the accuracy of the information contained in the resulting dataset, we compare its implied annual aggregate output, employment, total wage bill and goods exports with the official publicly available data. Figure C.3 in Appendix C shows that our dataset tracks well the evolution over time of these aggregates. Due to the reporting thresholds described above, aggregate exports in our sample fall a bit short of aggregate exports in the Customs data, but the gap is similar in the boom and bust periods (the average coverage is 91.8% in 2000-08 and 91.3% in 2009-13).[14]

We complement the firm-level data described above with yearly municipality-level data on the stock of vehicles per capita. The information on the stock of vehicles by municipality is provided by the Spanish Registry of Motor Vehicles, compiled by the General Directorate of Traffic (*Dirección General de Tráfico*), while the information on the population by municipality is provided by the Spanish National Statistical Office (*Instituto Nacional de Estadística*). When matching this municipality-level data with our firm-level data, we need to deal with the fact that the information on the location of firms is provided at the zip code level, and that the mapping between municipalities and zip codes is not one-to-one: larger municipalities are often assigned multiple zip codes and, in a very small number of cases, a single zip code is assigned to more than one municipality. In the former case, we associate the same value of the stock of vehicles per capita to all firms located in the same municipality, independently of the zip code of location; for firms in zip codes containing multiple municipalities, we associate with them a stock of vehicles per capita constructed as an average of the stocks of vehicles per capita across these municipalities.

We also employ information from two datasets provided to us by the Spanish Tax Agency (*Agencia Estatal de Administración Tributaria*, AEAT): first, aggregate data on municipality-to-municipality flows for *all* firms in the manufacturing sector, excluding sales of entities in the auto industry; and, second, firm-to-municipality sales only for those manufacturing firms in our sample that exported in the boom as well as in the bust.[15] The AEAT was willing to share with us only one year of data, so we work with data for the year 2006 as it is the first year for which a comprehensive digitization of the data is available.

When exploring the robustness of our results, we use information on additional variables. The underlying sources for these variables are discussed in Appendix B.

4 Identification Approach

In this section, we first describe our identification approach, and later highlight various potential threats affecting this strategy and how we seek to address them.

[14] Most of the gap in coverage is explained by the fact that a nontrivial share of Spanish exports is carried out by legal entities or individuals that are not registered as firms undertaking economic activity in Spain, and are thus exempted from submitting their financial statements to the Commercial Registry. The share of goods exports by non-registered entities was on average around 8% in 2010-13 (own calculations based on public Customs data).

[15] We thank Francesco Serti for having brought to our attention the existence of these data.

4.1 Geography-Based Proxies of Demand Changes

As explained in section 3.1, a key characteristic of the Great Recession in Spain is that it affected different regions differently. Panel (a) in Figure 3 illustrates this fact. The figure plots the standardized percentage change in domestic sales for the average manufacturing firm located in each of the 47 Spanish peninsular provinces and operating in at least one year of the boom period (2002-08) and at least one year of the bust period (2009-13).[16] The provinces where the average firm experienced a reduction in domestic sales smaller than the national average are in darker color, while those where the average firm experienced a larger reduction in domestic sales are in lighter color. Figure 3 thus illustrates that firms located in the Northern and Western regions saw changes in domestic sales larger (less negative) than the average, while firms located in the center of the country and in Southern and Eastern regions experienced relatively large domestic sales reductions. Furthermore, deviations from the national average are sizable in many cases.

The heterogeneity in the changes in domestic sales that we document in panel (a) of Figure 3 could have been caused by heterogeneity in supply factors or by heterogeneity in factors affecting local demand for manufacturing goods. We next propose an approach to measuring variation in local demand for manufacturing goods.

Our approach consists in proxying changes in local demand for manufacturing goods using observed changes in demand per capita for one particular type of manufacturing products (vehicles) for which we have highly geographically disaggregated data. Panel (b) in Figure 3 shows that there is substantial variation in the degree to which the number of vehicles per capita changed across provinces between the boom and the bust years.[17] Specifically, the provinces in the Northwest and in the Southwest experienced a relative increase in the number of vehicles per capita, while the region around Madrid and the provinces in the Northeast and along the Mediterranean coast experienced a relative reduction. As in panel (a) of Figure 3, the regional deviations from the national averages in panel (b) are large for many provinces.

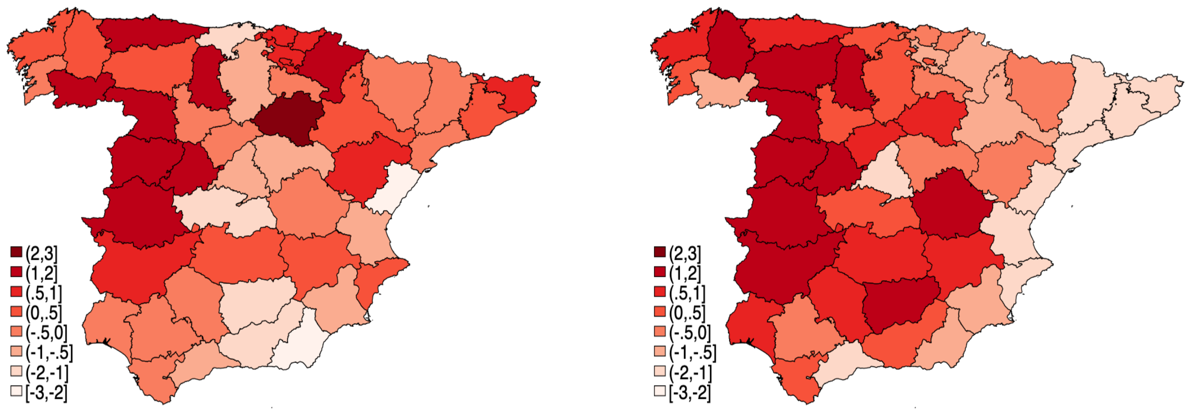
By illustrating provincial averages, Figure 3 hides substantial spatial variation at the sub-province level (across municipalities) in the boom-to-bust changes in both average firm-level domestic sales and changes in the stock of vehicles per capita. We illustrate this variation in Figure C.4 in Appendix C.4 for the case of the two most populated Spanish provinces: Madrid and Barcelona.

Our core empirical strategy exploits the variation illustrated in Figure C.4 to identify the impact of domestic demand shocks on firms' exports operating through its effect on the firms' domestic (Spain-wide) sales. Specifically, we divide our sample into a "boom" period (2002-08) and a "bust" period (2009-13), and assess the extent to which a demand-driven decline in a firm's domestic sales in the bust period relative to the boom period is associated with a relative increase in its export

[16] Figure C.2 in Appendix C.2 shows the yearly average number of firms and exporters by province for the period 2002-08. Economic activity in Spain is concentrated mostly in the coast (Galicia, Basque Country, Catalonia, Valencian Community, Murcia and Andalusia) and in the center (Madrid). Exporting firms are concentrated in the center (Madrid) and in the Mediterranean coast (Catalonia and Valencian Community).

[17] Changes in the stock of vehicles per capita between the boom and the bust years could have been due either to purchases of new vehicles or to scrapping of old ones. We measure the change in the stock, rather than just new purchases, to avoid that our measure of domestic demand for manufacturing firms is contaminated by the effect of the "cash for clunkers" program (*Plan PIVE*) that the Spanish government put in place during the bust period.

Figure 3: The Great Recession in Spain: Variation Across Provinces



(a) Relative Change in Domestic Sales (b) Relative Change in Vehicles per Capita

Notes: For each province, panel (a) illustrates the standardized percentage change in average firm-level domestic sales between the period 2002-08 and the period 2009-13, where the average is computed across manufacturing firms active in at least one year in both periods. Therefore, if this variable takes value p for a given province, it means that the average firm located in this province experienced a relative change in average yearly domestic sales between 2002-08 and 2009-13 that was p standard deviations above the change experienced by the average province. Panel (b) illustrates the standardized percentage change in vehicles per capita between the period 2002-08 and the period 2009-13. The color scale is in standard deviations as in panel (a).

sales between these two periods. We choose this 'long-differences' approach as our baseline because the macroeconomic evidence in Figure C.3 cleanly identifies the year 2009 as the break between two distinct periods. Having said this, we will show that our results are similar when breaking the sample into shorter subperiods, with the exception of panel regressions with yearly data, a frequency at which our instruments cease to have significant predictive power (see section 5.2).

To build a measure of the boom-to-bust change in domestic demand for each Spanish firm, we follow a two-step procedure. First, we use observed boom-to-bust changes in the stock of vehicles per capita at the municipality level as a proxy for the boom-to-bust changes in the demand for manufacturing goods in those municipalities. For this, we rely on a body of work (see Stock and Watson, 1999) documenting that durable goods consumption, and vehicle purchases in particular, are strongly procyclical, and are thus a useful proxy for changes in 'local demand'.[18] Second, with this measure of local demand at hand, we construct instruments capturing the boom-to-bust changes in domestic demand experienced by firms located in different Spanish municipalities. We do this in two distinct and complementary ways.

Our first instrument (or *local* instrument) builds on the work of Hillberry and Hummels (2008) highlighting the extremely localized nature of manufacturing shipments in the U.S.: we posit and verify that changes in municipality-level demand (as captured by changes in vehicles per capita) are a good predictor for changes in domestic (Spain-wide) sales of Spanish firms producing in the corresponding municipality.

Our second instrument (or *gravity-based* instrument) instead acknowledges that the majority of

[18]Relatedly, Mian et al. (2013) document that variation in the extent to which the U.S. subprime mortgage default crisis of 2007-10 affected household housing wealth in different areas in the U.S. translated into geographical variation in vehicle purchases. Hausman et al. (2019) and Waugh (2019) provide corroborating evidence in other settings.

a firm's sales are actually shipped outside its municipality of location, and constructs a theoretically-grounded measure that takes into account the exposure of firms to local demand shocks in municipalities *other* than their municipality of location. To measure the different exposure of each firm to other municipalities' local demand shocks, we rely on information on the location of these firms together with municipality-to-municipality trade flows data (aggregated across all manufacturing firms) for the year 2006. We use these data to estimate municipality-to-municipality gravity regressions, and use the estimated coefficients for log population and log distance to build the relevant weights needed to construct our gravity-based measure of firm-level exposure to local demand shocks, which is a weighted sum of the local shocks (i.e., changes in the stock of vehicles per capita) in all locations other than the one in which the firm is based. As a robustness check, we also experiment with alternative *gravity-based* instruments that rely on alternative sets of weights.

To be more precise, our two baseline instruments take the form

$$\Delta \ln Z_{id} = \begin{cases} \Delta \ln V_{\ell(i)} & \text{for the } \textit{local} \text{ instrument,} \\ \Delta \ln \left(\sum_{\ell' \neq \ell(i)} \left(Pop_{\ell'}\right)^{\hat{\beta}_{pop}} \left(Dist_{\ell'\ell(i)}\right)^{\widehat{\beta}_{dist}} V_{\ell'} \right) & \text{for the } \textit{gravity-based} \text{ instrument,} \end{cases} \tag{13}$$

where V_ℓ are vehicles per capita in municipality ℓ, $\ell(i)$ is the municipality in which firm i is located, $Pop_{\ell'}$ is population in location ℓ', $Dist_{\ell'\ell(i)}$ is the distance between municipalities ℓ' and $\ell(i)$, and $\widehat{\beta}_{pop}$ and $\widehat{\beta}_{dist}$ are estimates of the coefficients on (destination) population and on bilateral distance in a gravity equation estimated using Spain's municipality-to-municipality sales data. Our gravity-based instrument is closely related to the so-called Harris market access measure, a widely used measure of demand in economic geography studies (see Redding and Venables, 2004). Such a demand measure arises naturally in multi-market versions of the Melitz (2003) model; in fact, as Jacks and Novy (2018) show, this demand term also arises in more general frameworks, including all models that yield a structural gravity equation à la Anderson and van Wincoop (2003).[19]

Estimates of our municipality-to-municipality gravity equation for Spanish manufacturing flows in the year 2006 are presented in columns 1 through 3 of Table 1. These estimates are based on the municipality-to-municipality aggregate trade flows data provided to us by the Spanish Tax Agency (see section 3.3). The first column presents estimates for a specification with municipality of origin fixed effects, log population of the destination municipality, and log distance between origin and destination. The results illustrate the relevance of gravity forces, with shipments increasing in destination population with an elasticity of $\hat{\beta}_{pop} = 0.493$, and declining in distance with an elasticity of $\widehat{\beta}_{dist} = -0.429$. The inclusion in column 2 of dummies for own-municipality and own-province

[19] There are two important differences between our measure of market access and standard uses of market access as a proxy for demand. First, we measure the change in demand in a location not as the change in a destination fixed effect estimated from a gravity equation, but as the change in the stock of vehicles per capita in that location. Second, as the value of our instrument for a municipality excludes the change in vehicles per capita in such municipality, our instrument excludes what Redding and Venables (2004) denote the domestic market access term. The first difference is a consequence of the fact that we observe within-Spain flows only for the year 2006; the second one is due to our aim to make this second instrument distinct from the first one based on local demand (and also more plausibly exogenous). Similar market access terms have been defined in Harris (1954), Hanson (2005) and Donaldson and Hornbeck (2016).

Table 1: Estimates from Gravity Equations at Municipal Level

Dependent Variable:	Ln(Trade Flows)				
	(1)	(2)	(3)	(4)	(5)
Ln(Population)	0.493[a]	0.490[a]	0.485[a]	0.300[a]	0.322[a]
	(0.031)	(0.031)	(0.030)	(0.012)	(0.015)
Ln(Distance)	-0.429[a]	-0.378[a]		-0.150[a]	-0.145[a]
	(0.011)	(0.019)		(0.021)	(0.019)
Dummy for own-municipality flows		1.607[a]			
		(0.111)			
Dummy for own-province flows		0.131[b]			
		(0.065)			
Dummy for distance 5-10Km			-1.159[a]		
			(0.079)		
Dummy for distance 10-50Km			-1.924[a]		
			(0.090)		
Dummy for distance 50-100Km			-2.465[a]		
			(0.114)		
Dummy for distance 100-200Km			-2.718[a]		
			(0.113)		
Dummy for distance 200-500Km			-2.962[a]		
			(0.120)		
Dummy for distance 500-1000Km			-3.237[a]		
			(0.120)		
Dummy for distance >1000Km			-3.590[a]		
			(0.141)		
Observations	417,936	417,936	417,936	675,715	675,589
R-squared	0.30	0.31	0.30	0.15	0.31
Municipality of origin FE	Yes	Yes	Yes	Yes	No
Sector FE	No	No	No	Yes	No
Firm FE	No	No	No	No	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 municipality-level trade flows for manufacturing firms is for the 2006 fiscal year. Ln(Population) denotes the log of the population of the destination municipality in 2006. Ln(Distance) denotes the log of the distance, in kilometers, between the two municipalities in each pair. The estimates in columns 1 to 3 use municipality-to-municipality sales data; the estimates in columns 4 and 5 use firm-to-municipality data.

flows slightly reduces this distance elasticity, while these two dummies appear to have themselves a positive and significant effect on shipments. This suggests that part of the negative effect of distance on within-Spain municipality-to-municipality shipments is related to a discontinuous fall in shipments at the municipality border and at the province border. The extent of "home bias" at the municipality level is remarkably large: it implies that, *ceteris paribus*, shipments are $\exp(1.607) \approx 5$ times larger within a municipality than outside. The existence of such strong local home bias is in line with the findings of Hillbery and Hummels (2008) for the U.S., although the magnitude is larger in our setting.[20] This finding buttresses the potential relevance of our first *local* instrument.

[20] Díaz-Lanchas et al. (2019) also estimate a very sizeable 'zip code effect' on the basis of a random sample of shipments by road within Spain during the period 2003-07 (the C-Intereg database). Although we do not have access

Column 3 presents estimates for a specification analogous to that in column 1 but employing a set of distance dummies to capture the effect of distance on sales. The results show that, controlling for municipality of origin fixed effects and for the population of the municipality of destination, sales decay monotonically with distance.

In columns 4 and 5 of Table 1, we estimate gravity equations at the firm-destination level exploiting the information on the second of the datasets provided to us by the Spanish Tax Agency, which contains firm-to-destination shipments for the subset of firms that exported both in the boom as well as in the bust (see section 3.3). The specification in column 4 is analogous to that in column 1, while that in column 5 additionally accounts for firm fixed effects. The results are in line with those in column 1, but with somewhat smaller log population and log distance coefficients, as one would expect given that these specifications do not account for extensive margin variation in the set of municipalities firms sell to.[21]

4.2 Threats to Validity of the Identification Approach

The main concern with our identification approach is that our municipality-level measures of demand changes between the boom and the bust may be correlated with changes in marginal cost shifters affecting the firms located either in the corresponding municipalities or in neighboring ones. This exclusion restriction is central to the validity of our strategy, so we next outline how we try to deal with this potential threat to the validity of our identification approach.

First, we control in our specifications for sector fixed effects. Thus, we base our identification on observing how domestic sales and exports changed between the boom and the bust for different firms operating in the same sector but located in regions experiencing different exposure to local demand changes. By controlling for sector fixed effects, we control for sector-specific foreign demand shocks, sector-specific trade cost shocks, and domestic supply shocks affecting all Spanish firms operating in the same sector, irrespective of their location (see the definition of γ_{sx} in equation (6)). For example, these fixed effects control for shocks such as the expiration of the Multi Fiber Arrangement on January 1, 2005, which eliminated all European Union quotas for textiles imported from China, and which increased the competition that Spanish textile manufacturers faced both in the domestic and foreign markets.[22]

Second, as different firms operating in the same sector may experience different supply shocks,

to this information at the zip code level, we have obtained aggregated province-to-province shipments from that database. In Appendix D.4, we compare some aggregate statistics on firms' within-Spain sales from our sales data based on tax records and from the C-Intereg data. The extent of provincial home bias is very similar in both datasets.

[21] We have also estimated gravity equations that use the firm-to-destination data and include the border and distance dummies introduced in columns 2 and 3. Notably, we again find a remarkably large coefficient of 1.308 for the own-municipality dummy variable.

[22] Sector fixed effects may not effectively control for all heterogeneity across firms in their export demand shocks; specifically, firms located in different Spanish regions may be differentially affected by export demand shocks even if they operate in the same sector. A possible source of this heterogeneity in demand shocks is the different exposure of firms located in different Spanish regions to changes in demand in different foreign countries (e.g., firms located in Southern regions are more exposed to demand changes in Northern African countries than firms located in the north of Spain). In Appendix E.4, we provide suggestive evidence that this type of heterogeneity in export demand shocks does not, in practice, affect the validity of our instrument.

we also control throughout for firm-specific measures of productivity and labor costs. By controlling for changes in wages and productivity at the firm level, we aim to identify the effect that changes in local demand had on firms' exports through channels other than the internal devaluation channel. More specifically, these controls help address the concern that the reduction in unit labor costs observed in Spain during the period 2009-13 might have been heterogeneous across different Spanish municipalities in a manner that is correlated with our instruments.

A third concern is that, even after controlling for sector fixed effects and proxies for firm-level productivity and wage costs, there may still be unobserved *residual* location-specific marginal cost shifters that might be correlated with our proxy for changes in local demand. For instance, changes in labor payments not properly controlled for by our measure of firm-level wages may impact the purchasing power of consumers living in the corresponding municipalities, and thus affect vehicle purchases. This identification threat is likely to be particularly salient for our 'local' instrument. In the construction of our gravity-based instrument, we do not use information on the change in the number of vehicles per capita in the municipality of location of a firm, which helps assuage this concern as long as there is no a strong spatial correlation in the unobserved residual marginal cost shifters. Furthermore, to address concerns motivated by this possible spatial correlation in residual supply shocks, we present in section 6 estimates from regression specifications in which we control for additional measures of municipality-specific boom-to-bust changes in economic conditions.

A fourth concern relates to the presence of a nontrivial number of car manufacturers in our sample of Spanish manufacturing firms. These firms' supply shocks are especially likely to have impacted the boom-to-bust changes in the stock of vehicles per capita in their own municipality and in geographically close ones. More specifically, if a disproportionate share of cars in Spain was sold in municipalities that are geographically close to where the car was manufactured, supply shocks in these firms may affect the number of vehicles per capita not just by affecting the purchasing power of consumers in certain municipalities, but by affecting directly the supply of cars in those municipalities. Roughly three quarters of all cars purchased in Spain are imported (as indicated by data from the Spanish National Institute of Statistics); thus, supply shocks affecting car manufacturers are likely to have a limited impact on the total amount of cars in Spain. However, to deal with this threat to identification, we exclude all firms operating in the auto industry (NACE Rev. 2 code 29) in all the regressions we present. Additionally, we also explore in section 6 how our results are impacted when excluding from our sample: (i) all firms located in a zip code that hosts at least one firm in the auto industry employing more than 20 workers; (b) all firms located in a zip code or in the proximity of a zip code with a significant share of manufacturing employment accounted for by the auto industry; and (c) all firms producing in sectors that are either leading input providers or leading buying industries of the vehicles manufacturing industry.

Finally, it is important to remark that, as illustrated in equation (12), any unobserved factor costs that are negatively correlated with either of the two instruments we use will cause the corresponding IV estimator to be positively biased. For example, if the tightening of the credit supply in a region caused firms' marginal production costs to increase and consumers' demand to fall, the resulting endogeneity of our instrument would bias our IV estimator upwards. Thus, negative

IV estimates of the elasticity of firm-level exports with respect to a firm's domestic sales would still reflect patterns in the data that would be inconsistent with the constant marginal cost model described in section 2, and that would be suggestive of the existence of a negative relationship between demand-driven changes in domestic sales and exports.

5 Baseline Results

In this section, we present our baseline results on the impact of demand-driven changes in domestic sales on firms' behavior in the export market. Specifically, we present in sections 5.1 and 5.2 evidence of the impact of the Great Recession on Spanish firms' intensive margin of exports. In section 5.3, we present analogous evidence of its impact on the extensive margin.

5.1 Intensive Margin

Table 2 presents OLS estimates of the elasticity of boom-to-bust changes in firms' exports with respect to boom-to-bust changes in domestic sales for continuing exporters – i.e., firms that exported both in the boom as well as in the bust. There are $8,009$ such firms in our dataset.

As discussed in section 2, unobserved (residual) supply factors tend to make the OLS estimate of a firm's change in domestic sales on its change in foreign sales positive even in a world with constant marginal costs. Conversely, measurement error in both total sales and exports tends to make this OLS estimate negative. As illustrated in column 1 of Table 2, when no controls are included, we estimate an OLS elasticity of export flows with respect to domestic sales that is very close to zero. In column 2, we control for the change in firms' productivity (estimated following the procedure in Gandhi et al., 2020, as detailed in Appendix F), and in column 3 for the change in the firm's average wages. Consistent with the discussion in section 2, controlling for these supply shocks reduces the OLS estimate of the coefficient on domestic sales. In fact, the coefficient turns significantly negative (-0.298), indicating that, once we control for the observable part of firms' supply shocks, domestic sales and exports are negatively correlated. Columns 4, 5 and 6 aim to control for additional unobserved determinants of firms' marginal costs that vary between the boom and the bust. To do so, and motivated by the specification in equation (10), we sequentially add sector fixed effects (in column 4), province fixed effects (column 5) and municipality fixed effects (column 6). The resulting estimates continue to be negative and indicate that a 1% decrease in a firm's domestic sales, keeping its productivity and average wages constant, is associated with close to a 0.3% increase in its overall export flows.

In Table 3, we turn to our baseline two-stage least squares (TSLS) estimates of the elasticity of the firm's boom-to-bust change in exports with respect to its boom-to-bust demand-driven change in domestic sales. As discussed in section 2, if our local and gravity-based instruments are orthogonal to both unobserved supply factors and the measurement error in both total sales and exports, and firms' marginal costs are constant, the corresponding TSLS estimators should converge to zero. The first-stage estimates reported in columns 1 to 4 of Table 3 reveal that firms located in municipalities that experienced a larger drop in either their local (panel A) or gravity-based

Table 2: Intensive Margin: Ordinary Least Squares Estimates

Dependent Variable:	ΔLn(Exports)					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Domestic Sales)	0.063	-0.209[a]	-0.298[a]	-0.292[a]	-0.284[a]	-0.271[a]
	(0.044)	(0.049)	(0.043)	(0.032)	(0.032)	(0.036)
ΔLn(TFP)		1.142[a]	1.448[a]	1.535[a]	1.522[a]	1.514[a]
		(0.043)	(0.046)	(0.057)	(0.055)	(0.057)
ΔLn(Average Wages)			-0.744[a]	-0.723[a]	-0.712[a]	-0.706[a]
			(0.062)	(0.072)	(0.070)	(0.067)
Observations	8,009	8,009	8,009	8,009	8,009	7,502
R-squared	0.001	0.100	0.126	0.162	0.171	0.278
Sector FE	No	No	No	Yes	Yes	Yes
Province FE	No	No	No	No	Yes	No
Municipality FE	No	No	No	No	No	Yes

Notes: [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. The estimation sample includes all firms exporting in at least one year in the period 2002-2008 and in the period 2009-2013.

(panel B) proxy for demand also suffered a larger decline in their domestic (Spain-wide) sales. This relationship is robust to controlling for our measures of firms' changes in productivity and labor costs, for sector fixed effects and, when using the local instrument, for province fixed effects: the statistics of F-tests of the null hypothesis that the coefficient on our instruments is equal to zero in the first-stage regressions is in all specifications above threshold values generally applied to detect weak instrument problems, the only exception being the value of 7.85 in column 1 of panel B.[23]

The second-stage estimates (reported in columns 5 to 8 of Table 3) reveal elasticities of exports with respect to domestic sales that are significantly larger in absolute value than the OLS elasticities reported in Table 2.[24] This suggests that, even after controlling for sector fixed effects and firm proxies of productivity and average labor costs, there still remains substantial unobserved determinants of firms' marginal costs that induce a spurious positive correlation between their sales in the domestic and foreign markets. Our preferred estimates in column 8 indicate an elasticity of exports with respect to domestic sales of around -1.3 for our local instrument and of around -1.6 for our gravity-based instrument. These negative estimates are suggestive of the firm's marginal cost function not being flat.

One might be concerned that, because firms' total sales are a key input in the computation of our proxy for a firm's productivity, our empirical results are just unveiling a mechanical negative correlation between exports and domestic sales once one holds total sales revenue constant (by controlling for it). Although log TFP and log total sales are obviously positively correlated (as one would expect in light of the model described in section 2), the correlation is far from perfect,

[23]In Appendix Figure C.5, we present binned scatterplots of both the first-stage and reduced-form relationships.

[24]The unrealistically high point estimate and the corresponding large standard error reported in column 5 of panel B should be discounted on the basis that, as shown in column 1, our instrument is weak in this regression specification.

Table 3: Intensive Margin: Two-Stage Least Squares Estimates

	Panel A: Local Instrument							
Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔLn(Domestic Sales)					-2.428^a	-1.272^a	-1.185^a	-1.327^a
					(0.685)	(0.306)	(0.276)	(0.337)
ΔLn(Vehicles p.c. in municipality)	0.310^a	0.455^a	0.474^a	0.418^a				
	(0.064)	(0.071)	(0.072)	(0.073)				
ΔLn(TFP)		0.800^a	0.992^a	0.980^a		1.983^a	2.320^a	2.533^a
		(0.027)	(0.031)	(0.032)		(0.244)	(0.277)	(0.323)
ΔLn(Average Wages)			-0.606^a	-0.511^a			-1.279^a	-1.240^a
			(0.035)	(0.043)			(0.184)	(0.178)
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
Province FE	No	No	No	Yes	No	No	No	Yes
F-statistic	23.19	41.64	43.26	33.10				

	Panel B: Gravity-based Instrument							
Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔLn(Domestic Sales)					-10.068^a	-2.081^a	-1.751^a	-1.607^a
					(3.454)	(0.319)	(0.238)	(0.248)
ΔLn(Dist-Pop-Weighted Vehicles p.c.)	0.339^a	1.194^a	1.346^a	1.312^a				
	(0.121)	(0.145)	(0.135)	(0.119)				
ΔLn(TFP)		0.829^a	1.031^a	1.023^a		2.623^a	2.876^a	2.810^a
		(0.028)	(0.029)	(0.028)		(0.241)	(0.222)	(0.213)
ΔLn(Average Wages)			-0.621^a	-0.526^a			-1.620^a	-1.387^a
			(0.037)	(0.047)			(0.174)	(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	7.85	67.47	99.84	122.44				

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors, clustered by municipality in panel A and by province in panel B, 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. Vehicles p.c. denotes the stock of vehicles per capita in a municipality. 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 1 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 the instrumental variable equals zero.

particularly when considering log changes in these variables. More specifically, the correlation between log changes in our proxy for a firm's productivity and log changes in observed total sales is 0.31 at the yearly level, and is 0.56 when looking at boom-to-bust 'long differences' in these variables. To further assuage this concern, we explore in section 6.3 the robustness of our results to using an alternative measure of firms' productivity; this alternative measure uses information on firms' value-added instead of total sales, and has a much lower correlation with this latter variable.

In terms of the quantitative relevance of our results, it is worth emphasizing that an elasticity of -1.6 does not necessarily imply a more-than-complete substitution of exports for domestic sales. For a firm with an initial export share of $100 \times \chi\%$, a demand-driven drop of €100 in their domestic sales would lead to a €160 $\times$ $(\chi/(1-\chi))$ increase in exports. For example, for every €100 of lost domestic sales, a firm with an export share of 25% would able to recoup €53.3 via exports, while a firm with an export share of one-third would be able to recoup €80.[25]

In terms of the statistical significance of our results, it is important to remark that, unless otherwise noted, standard errors are clustered by municipality when using the local instrument and by province when using the gravity-based instrument. It is also worth highlighting that none of our instruments fall in the class of shift-share instruments studied recently by Adão et al. (2019), Borusyak et al. (2020) and Goldsmith-Pinkham et al. (2020), among others. The reason is that our instrument does not use changes in a weighted-average of vehicles per capita across municipalities but log changes in such weighted average.[26] Although our choice of functional form prevents us from computing the standard errors according to the formulas introduced in Adão et al. (2019), it is conceivable that our standard error estimates suffer from the downward bias that typically affect clustered standard errors when the instrument is of the shift-share type. We revisit this question in section 6.2, where we provide suggestive evidence showing that the bias affecting our standard errors, if present, is likely to be very small.

5.2 Panel Specifications

Although the evidence shown in Figures C.1 and C.3 in Appendix C shows that, from a macroeconomic perspective, there are clearly two distinct periods (a boom and a bust period) in our sample, our long-differences approach comparing the boom to the bust has the limitation that it does not allow to control for location-specific trends that could help account for the evolution of location-specific unobserved supply conditions. To address this limitation, we report in Table 4 TSLS estimates for various specifications that exploit the higher frequency of our data, and that thus allow for the inclusion of municipality-specific time trends in our estimating equation.

In columns 1 to 3 of Table 4, we present estimates from panel specifications in which each observation in the time dimension corresponds to a three-year rolling average, starting with 2002-04 all the way to 2011-13. As in Table 2, we report results for our local instrument in panel A, and for our gravity-based instrument in panel B. The OLS, first-stage, and second-stage results are all qualitatively similar to those reported in Tables 2 and 3, although the key second-stage elasticities are slightly larger in this case (-1.4 and -2.2, respectively). In columns 4 to 6, we present analogous results using two- (rather than three-year) rolling averages. The results are again qualitatively similar, but our instruments become significantly weaker in this case, especially the gravity-based one. Finally, we report in columns 7 to 9 results in which each observation in the

[25] The median export share among the 8,009 firms exporting in both boom and bust periods is 16.2%.

[26] The reason for employing log changes to build this instrument is that, according to the multi-destination model in Appendix E.2, log changes in domestic sales are linearly related to log changes in a weighted average of local demand shifters across the different domestic markets.

Table 4: Panel Regressions

	Panel A: Local Instrument								
Data Frequency:	3-year Rolling Average			2-year Rolling Average			Annual Data		
	OLS	1st Stage	2SLS	OLS	1st Stage	2SLS	OLS	1st Stage	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Ln(Domestic Sales)	-0.259^a		-1.441^a	-0.259^a		-1.040^c	-0.273^a		2.222
	(0.019)		(0.528)	(0.017)		(0.631)	(0.014)		(5.777)
Ln(Vehicles p.c.) in municipality)		0.278^a			0.193^a			0.023	
		(0.056)			(0.051)			(0.050)	
Ln(TFP)	1.420^a	0.930^a	2.517^a	1.407^a	0.920^a	2.125^a	1.385^a	0.920^a	-0.911
	(0.045)	(0.027)	(0.486)	(0.043)	(0.028)	(0.582)	(0.042)	(0.030)	(5.313)
Ln(Average Wages)	-0.629^a	-0.416^a	-1.119^a	-0.614^a	-0.394^a	-0.921^a	-0.581^a	-0.373^a	0.349
	(0.044)	(0.031)	(0.229)	(0.040)	(0.030)	(0.256)	(0.035)	(0.029)	(2.152)
Observations	66,711	66,710	66,710	65,709	65,708	65,708	60,199	60,198	60,198
F-statistic		24.45			14.35			0.21	
	Panel B: Gravity-based Instrument								
Data Frequency:	3-year Rolling Average			2-year Rolling Average			Annual Data		
	OLS	1st Stage	2SLS	OLS	1st Stage	2SLS	OLS	1st Stage	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Ln(Domestic Sales)	-0.259^a		-2.159^a	-0.259^a		-2.380^a	-0.274^a		2.066
	(0.022)		(0.242)	(0.017)		(0.603)	(0.018)		(1.580)
Ln(Dist-Pop-Weighted Vehicles p.c.)		1.016^a			0.775^a			0.677	
		(0.157)			(0.279)			(0.453)	
Ln(TFP)	1.420^a	0.941^a	3.184^a	1.407^a	0.923^a	3.359^a	1.382^a	0.921^a	-0.773
	(0.052)	(0.021)	(0.226)	(0.055)	(0.023)	(0.549)	(0.064)	(0.027)	(1.481)
Ln(Average Wages)	-0.629^a	-0.422^a	-1.417^a	-0.614^a	-0.395^a	-1.449^a	-0.579^a	-0.373^a	0.295
	(0.037)	(0.032)	(0.136)	(0.034)	(0.030)	(0.258)	(0.037)	(0.031)	(0.582)
Observations	66,711	66,711	66,711	65,709	65,709	65,709	60,199	60,199	60,199
F-statistic		41.73			7.70			2.24	

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 to 3 is constructed calculating three-year rolling 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 to 6, we calculate two-year rolling averages, where the periods are 2002-2003, 2003-2004, etc., for a total of 11 periods. In columns 7 to 9, we use the original annual data with 12 periods between 2002 and 2013.

time dimension corresponds to a year, exploiting thus the full time variation in our annual data. Although the OLS estimates remain stable, our instruments become very weak in this case.

There are several possible reasons for why spatial variation in vehicle sales fails to predict demand conditions at the yearly level. First, vehicle sales are not only pro-cyclical (as other durable goods are) but also tend to be lead indicators for recessions. For instance, in our own aggregate data from Spain, vehicles per capita declined already in 2008 (see Figure C.1), while the recession in Spain only unfolded in 2009. Second, the precise extent to which vehicles sales led or lagged the drop in demand across municipalities in Spain during the Great Recession varies across municipalities. Both of these concerns affect the relevance of the instruments when each period corresponds to a year, but they do not do so when each observation in the time dimension corresponds to an average across years, as in our baseline boom-to-bust regressions and in the three-and two-year rolling window specifications.[27]

5.3 Extensive Margin

To study the impact of demand shocks on the extensive margin of exporting, we again divide the sample period into a boom (2002-08) and a bust period (2009-13), and explore how demand-driven changes in domestic sales impact changes in firms' probability of exporting between these two periods. More specifically, we use data on all firms in our sample that are active in the domestic market in both the boom and the bust, and compute TSLS estimates of a linear probability model in which a firm's dummy capturing positive exports in a given period (boom or bust) is regressed on firm, sector-period and province-period fixed effects, and the log of the firm's average wages and productivity measures, with log domestic sales in a given period instrumented by either the local or the gravity-based proxy for the firm's demand in that period.[28] We present in Table 5 estimates that use the gravity-based instrument, and in Appendix G.1 estimates that rely on the local instrument. In addition, for both types of instruments, we also present estimates for specifications in which the dependent variable is the proportion of years in a given period (boom or bust) that a firm exports.

Column 1 in Table 5 reports the first-stage estimates. As in Table 3, the results indicate that domestic sales fell more for firms located in municipalities that, according to our gravity-based instrument, experienced a larger drop in domestic demand. The F-stat (86.32) is, as in our intensive margin specifications, well above standard threshold values. Columns 2 and 3 present OLS and TSLS estimates of the link between domestic sales and export status, while columns 4 and 5 report OLS and TSLS estimates of the link between domestic sales and the proportion of years exported. The results of these two specifications deliver statistically significant estimates of opposite signs but, in both cases, these are quantitatively very small. As shown in Table G.1 of Appendix G.1, we find similarly weak and mixed extensive margin effects when using the local instrument.

Taken together, these results lead us to conclude that the vent-for-surplus mechanism did not appear to operate strongly via the extensive margin (i.e., via entry and exit from the export market). This result aligns with the fact, discussed in section 3.2, that more than 90% of the growth in Spanish exports during the bust period was explained by continuing exporters. There are two potential explanations that make our muted extensive margin results compatible with the sizeable intensive margin effects discussed in section 5.1. The first explanation relates to the fact that we only have data on aggregate exports, and thus changes in the extensive margin in our context refer to entry and exit from export markets *altogether*, which is a decision involving much larger investments than entry and exit from specific export markets. Second, a richer model than that described in section 2.1 could easily account for the muted impact of demand-driven changes in domestic sales on the extensive margin of trade. For instance, a model that incorporates sunk

[27]While the first of these two concerns could be addressed by including lags of the instruments in the first-stage specification, the second one is harder to address. In Table G.4 in Appendix G, we report results of yearly regressions that also include lags of the instruments in the first-stage specification. The instruments (regardless of whether it is the local or the gravity-based one) continue to be weak in this case.

[28]Our results are similar when we use as left-hand-side variable a dummy that treats a firm as an 'exporter' only if it exports for two or more years in a given period.

Table 5: Extensive Margin: Two-Least Squares Estimates

Dependent Variable:		Export Dummy		Proportion of Years	
	1st Stage	OLS	2SLS	OLS	2SLS
	(1)	(2)	(3)	(4)	(5)
Ln(Domestic Sales)		0.021^a	-0.099^a	0.008^b	0.040^b
		(0.005)	(0.034)	(0.004)	(0.019)
Ln(Dist-Pop-Weighted Vehicles p.c.)	1.024^a				
	(0.110)				
Ln(TFP)	1.169^a	0.068^a	0.204^a	0.062^a	0.024
	(0.018)	(0.007)	(0.039)	(0.005)	(0.020)
Ln(Average Wages)	-0.589^a	-0.046^a	-0.114^a	-0.041^a	-0.022^b
	(0.015)	(0.007)	(0.022)	(0.004)	(0.010)
Observations	125,054	125,054	125,054	125,054	125,054
F-statistic	86.32				

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province reported in parenthesis. Ln(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 1 of Table 1. F-statistic denotes the corresponding test statistic for the null hypothesis that the coefficient on Ln(Dist-Pop-Weighted Vehicles p.c.) equals zero. All specifications include firm and sector-period fixed effects. The estimation sample includes all firms selling in the domestic market in at least one year in the period 2002-08 and in the period 2009-13.

costs of exporting will tend to generate hysteresis in exporting status and, in this context, any shock to export profitability may have a very different impact on the intensive and extensive margins of trade depending on firms' expectations about its persistence (see Dickstein and Morales, 2018). In sum, the estimates in Table 5, and the fact that most of the growth in Spanish exports in the years following the Great Recession was due to firms that were already exporting during the boom, strongly suggest that the relationship between demand-driven changes in domestic sales and the extensive margin of exports in Spain was not quantitatively important in this period. Consequently, we focus in the remainder of this paper on variants of the intensive margin results in Table 3.

6 Robustness

In this section, we present additional evidence that further supports the empirical relevance of the "vent-for-surplus" mechanism. Specifically, we present estimates of regression specifications that address some specific sources of endogeneity that could bias the IV estimates presented in section 5. To save on space, we only report here regressions that use our gravity-based instrument, and present in Appendix G.1 analogous results that use our local instrument.

6.1 Further Purges of the Auto Industry

While the sample used to compute the estimates in Table 3 excludes firms classified in the manufacturing of motor vehicles sector, one may still be concerned that the salient presence of firms in

Table 6: 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 'neighboring' 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(Domestic Sales)		-1.693^a		-1.663^a		-1.693^a		-1.864^a
		(0.277)		(0.371)		(0.317)		(0.343)
ΔLn(Dist-Pop-Weighted Vehicles p.c.)	1.290^a		1.372^a		1.311^a		1.238^a	
	(0.119)		(0.166)		(0.132)		(0.130)	
ΔLn(Avg. Wages)	-0.511^a	-1.372^a	-0.505^a	-1.364^a	-0.490^a	-1.325^a	-0.491^a	-1.459^a
	(0.052)	(0.166)	(0.070)	(0.222)	(0.053)	(0.183)	(0.049)	(0.165)
ΔLn(TFP)	1.022^a	2.867^a	1.006^a	2.801^a	1.008^a	2.817^a	1.010^a	3.011^a
	(0.034)	(0.243)	(0.050)	(0.300)	(0.037)	(0.277)	(0.036)	(0.285)
Observations	7,180	7,180	4,595	4,595	6,131	6,131	6,072	6,072
F-statistic	118.30		68.58		98.78		91.09	

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.

that industry in a given municipality might lead to a negative association between the boom-to-bust changes in the stock of vehicles per capita and the boom-to-bust changes in *residual* marginal costs shifters of the firms located in that municipality (even if they operate in other industries). This would be the case if the boom-to-bust drop in the number of vehicles per capita in a municipality was caused by an exogenous increase in marginal costs affecting the firms in the motor vehicles industry, and this negative supply shock was transmitted to other firms within the same municipality, shifting upwards their marginal cost functions.[29] Notice however that this source of endogeneity in our instrument would cause the TSLS estimates presented in Table 3 to be upward biased, as unobserved shocks that increase firms' marginal costs would have a negative impact on their exports. In order to evaluate the robustness of our estimates to this concern, we report in Table 6 TSLS estimates for regressions specifications analogous to those in columns 4 and 8 of Table 3, but for four alternative samples. In panel A, we exclude from our sample all firms located in a zip code that ranks in the top 25% of zip codes by share of manufacturing employment accounted for by motor-vehicles producers (as computed from our micro-level data). In panel B, we further restrict the sample relative to panel A by excluding all firms located in a zip code in which at least one motor-vehicles producer with more than 20 workers operates. In panel C, we exclude all firms from zip codes neighboring a zip code that ranks in the top 25% of zip codes by share of manufacturing employment in motor-vehicles producers.[30] Finally, in panel D, we exclude all firms

[29] For example, the post-2009 trade collapse may have increased the input costs for firms in the motor vehicles industry, which may have passed these higher costs through to their buyers, both final consumers and other firms located in the same municipality but operating in different industries.

[30] We identify two zip codes as neighboring each other if they share the first four digits of their 5-digit code.

producing in sectors that are either one of the two top leading input providers or two top leading buying industries of the vehicles manufacturing industry. In all panels, we obtain slightly larger (in absolute value) estimated elasticities than in our baseline results, consistently with the hypothesis that these sample restrictions attenuate concerns about our estimates being upward biased.

6.2 Alternative Instruments

As described in detail in section 4.1, the value of our gravity-based instrument for each municipality is computed as the log change in a weighted average of vehicles per capita in every other municipality, where the weight attached to a municipality depends on its population and the distance to the municipality of origin, and where the elasticities of this weight with respect to these covariates correspond to the estimates reported in column 1 of Table 1. In columns 2 to 4 of Table 7, we test the robustness of our results to instruments constructed similarly to our baseline gravity-based instrument, but with weights that depend on distance and population in different ways. In particular, after reproducing our baselines estimates in column 1, columns 2 to 4 present results corresponding to gravity-based instruments built using weights that depend on distance and population according to the estimates reported in columns 2 to 4 of Table 1. As a reminder, the first of these differs from our baseline in that it includes own-municipality and own-province dummies, the second one uses a more flexible specification to estimate the impact of distance on municipality-to-municipality shipments, and the third one relies on a gravity equation analogous to the one we use to build our baseline instrument, but exploits firm-to-municipality shipment flows instead of municipality-to-municipality ones. The resulting second-stage elasticities of exports to demand-driven changes in domestic sales are in all three cases very similar to the baseline ones.[31]

In column 5 of Table 7, we present results based on instruments computed in an analogous manner as in our baseline specification, but with the difference being that, instead of using weights that only vary bilaterally between municipalities and are predicted by a gravity equation, we use weights that vary across firm-municipality pairs and that correspond to the *actual* domestic sales share of each firm in each Spanish municipality in 2006. As in our baseline instrument, we assign a zero weight to the municipality of location of each firm. This specification results in a larger (in absolute terms) elasticity of exports to domestic sales, although the difference from our baseline of -1.6 is approximately only one standard deviation.

To address potential concerns about downward bias in our gravity estimates due to the many zeroes in our sample (see Fitzgerald and Haller, 2018), we present in column 6 estimates that rely on an instrument that is analogous to our baseline except for relying on weights built as if we had obtained a coefficient of 1 on log population, a coefficient of -1 on log distance, and a coefficient of 1.5 on the province dummy as our gravity equation estimates. These are standard estimates in the intranational gravity literature that uses province-to-province trade flows.[32] The results are again

[31] The results are also virtually identical when using the estimated distance and population elasticities reported in column 5 (rather than column 4) of Table 1.

[32] See among, many others, Wolf (2000) for the U.S., Garmendia et al. (2012) for Spain, Helliwell (1996) for Canada, Combes et al. (2005) for France, Poncet (2005) and Xing and Li (2011) for China, Daumal and Zignago (2010) for Brazil, and Nitsch (2002) and Helble (2007) for Germany. We have run an analogous gravity equation by aggregating

Table 7: Alternative Instruments

	Panel A: First Stage with Alternative Instruments						
Dependent Variable:	ΔLn(Domestic Sales)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Gravity: mun-mun flows (Baseline)	1.312[a] (0.119)						
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Baseline incl. own mun. & prov. dummies		1.334[a] (0.119)					
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Gravity: distance dummies			0.954[a] (0.128)				
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Gravity: firm-mun flows				1.303[a] (0.106)			
ΔLn(Weighted Vehicles p.c.) Weights: firm-level mun. shares)					0.521[a] (0.111)		
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Fixed coefficients: $\beta_{pop} = 1$, $\beta_{dist} = -1$						0.405[a] (0.111)	
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Baseline in levels							0.967[a] (0.115)
Adão et al. (2019) std. error							(0.208)
Observations	8,009	8,009	8,009	7,906	7,850	8,009	8,009
F-statistic	122.44	126.55	55.90	151.03	21.82	13.21	70.94
							21.54
	Panel B: Second Stage with Alternative Instruments						
	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Domestic Sales)	-1.607[a]	-1.614[a]	-1.336[a]	-1.642[a]	-2.162[a]	-1.626[a]	-1.701[a]
	(0.248)	(0.245)	(0.228)	(0.214)	(0.412)	(0.418)	(0.262)
							(0.271)
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, we first report standard errors clustered by province, and below we report standard errors that apply the formula in Adão et al. (2019). All specifications include firm-level log differences in TFP and log differences in average wages as additional controls. Additionally, all specifications also include sector fixed effects. For a detailed description of each of the instruments, see text.

very similar to those in our baseline specification.

Finally, in column 7, we present results computed using an instrument that equals the exponential of our baseline instrument and that, consequently, is simply a weighted sum of the change in vehicles per capita across different municipalities. This instrument belongs to the category of shift-share instruments considered by Adão et al. (2019), thus allowing us to compute standard errors in the manner recommended in that paper. Given that both the baseline instrument and the shift-share instrument used in column 7 rely on the same identification assumptions, it is reassuring that the second-stage point estimates they yield are very similar (-1.607 vs. -1.701). To evaluate the possible bias of standard errors clustered by province in our empirical application, we compute in column 7 both standard errors clustered by province (first number in parenthesis) and the standard errors suggested in Adão et al. (2019) (second number in parenthesis). As illustrated

our AEAT data at the province level, and have found a coefficient of 1.332 on log population, -1.091 on log distance, and 1.449 on a dummy for own province flows (see Table G.14 in Appendix G.8).

in column 7 of Table 7, the downward bias affecting the standard errors based on clustering by province is very small; while the second-stage standard error that clusters by province is 0.262, the standard error computed according to the formula introduced in Adão et al. (2019) is 0.271.[33]

In section 6.5, we discuss additional alternative instrumentation strategies related to the deep roots of the differential fall in demand across Spanish regions.

6.3 Controlling for Additional Confounding Factors

The controls and various fixed effects included in our baseline specification may still not fully account for the impact on exports of certain marginal cost shifters that could be correlated with our instruments, thus biasing our estimates. We assess here the robustness of our results to the inclusion of additional proxies for firm- and municipality-level cost shifters.

The first set of additional controls aims to avoid possible biases in our IV estimates arising from the dual nature of the Spanish labor market, with large differences in dismissal costs between temporary- and permanent-contract workers. This difference in dismissal costs might have led firms to shed a disproportionate number of temporary workers during the bust. If so, given the average differences in skill and experience between both types of workers, firms shedding temporary workers may have undergone a skill- and experience-upgrading that changed firms' marginal production costs in a way that is not properly accounted for by our firm-level measures of productivity and average wages. The second set of additional controls aims to proxy for factor costs other than labor costs; more specifically, it aims to control for changes in the financial costs (explicit via interest rates, or implicit via rationing) that firms experienced during the Great Recession years.

After reproducing our baseline estimates in column 1, in column 2 of Table 8 we additionally control for the firm-level boom-to-bust change in the share of temporary workers. The negative and statistically significant point estimate indicates that firms that shed a disproportionate number of temporary workers during the bust period experienced a larger increase in exports, which is in line with our hypothesis above about the differences in productivity between temporary and permanent workers. The IV estimate of the causal effect of demand-driven changes in domestic sales on exports is however only slightly increased (elasticity of -1.639).[34] In columns 3 and 4, we introduce municipality-level controls for local labor market conditions. Column 3 includes the same change in the ratio of temporary workers over total employment as in column 2, but computed at the municipality level. In column 4, we further control for a municipality-level measure of the change in the manufacturing employment per capita. The inclusion of these two controls has a negligible impact on the main coefficient of interest, and only the second one has a significant effect

[33] The similarity in our application between the standard errors that cluster by province and those suggested in Adão et al. (2019) is hardly surprising: municipalities located in the same province assign similar weights to every other municipality and, thus, firms located in the same province are, according to our shift-share instrument, similarly exposed to changes in the stock of vehicles per capita across all Spanish municipalities.

[34] We obtain similar results when instead controlling for the (initial) firm-level share of temporary workers during the boom period. Specifically, firms that entered the bust period with a larger share of temporary workers (and thus had a larger potential to affect their skill composition when transitioning to the bust period) experienced higher export growth in the bust relative to the boom, but the estimate of our parameter of interest is largely unaffected.

Table 8: Confounding Factors

Dependent Variable:	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Domestic Sales)	-1.607^a	-1.639^a	-1.618^a	-1.632^a	-1.678^a	-1.680^a	-1.611^a
	(0.248)	(0.251)	(0.263)	(0.259)	(0.251)	(0.257)	(0.264)
ΔShare of Temp. Workers		-0.250^a					
(firm level)		(0.089)					
ΔShare of Temp. Workers			-0.019				
(muni. level)			(0.170)				
ΔManufacturing Empl. p.c.				-0.272^a			
(muni. level)				(0.050)			
ΔLn(Financial Costs)					-0.027^c		
(firm level)					(0.014)		
Financial Costs in Boom						-0.008	
(firm level)						(0.015)	
Financial Costs in Boom							-0.039
(muni. level)							(0.041)
Observations	8,009	7,640	7,743	7,745	6,879	6,945	7,741
F-Statistic	122.44	131.97	138.65	136.93	88.43	89.27	139.26

Note: ^{a} denotes 1% significance, ^{b} denotes 5% significance, ^{c} denotes 10% significance. Standard errors clustered by province reported in parentheses. In all specifications, ΔLn(Domestic Sales) is instrumented by ΔLn(Distance-Population-Weighted 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 fixed effects.

on exporting.[35]

In columns 5 to 7 of Table 8, we study potential confounding effects related to financial costs. We construct a measure of the financial costs that each firm faces in each period as 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 impact of this measure on firms' changes in exports is statistically different from zero only at the 10% level, and including this variable only has a marginal effect on the estimate of the elasticity of exports to domestic sales (which becomes -1.678). 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. Although we do not have measures of firms' credit applications and whether these were denied, one may conjecture that firms whose financial costs were larger in the boom were more likely to suffer credit rationing in the bust. Regardless of whether we measure financial costs in the boom using each firm's financial information (column 6) or as the average financial costs of all other firms located in the same municipality (column 7), our results indicate that either credit rationing had little impact on firms'

[35]Firms located in municipalities with larger declines in manufacturing employment per capita experienced higher export growth, potentially due to workers' extra effort in reaction to the reduction in employment opportunities in their municipality.

exports or our conjecture that it may be measured through the firms' financial costs in the boom has little empirical support.[36]

6.4 Alternative Productivity Estimates

We next test the robustness of our results to an alternative approach to measuring firms' productivity. Columns 1 and 2 in Table 9 replicate our baseline OLS and IV estimates presented in column 4 of Table 2 and column 8 of Table 3, while columns 3 and 4 of Table 9 present estimates of specifications that differ exclusively on the productivity measure.

Consistently with the model described in section 2, our baseline and alternative productivity measures exploit the assumption that firms: (a) face a CES demand function and are monopolistically competitive in both the domestic and the foreign market; (b) take all factor prices as given. The two approaches we implement differ however on the assumptions we impose on the shape of the production function and on the underlying output series they employ in the estimation. In both estimation approaches, which we describe in detail in Appendix F, we invoke optimality conditions for the static inputs (labor and materials) in order to estimate the relevant parameters of the production function and, in this sense, both approaches are specific cases of the general estimation framework in Gandhi et al. (2020). But while our baseline approach exploits data on the firm's total sales, the alternative approach in columns 3 and 4 uses information on the firm's value added.

As anticipated in section 5.1, a general concern with our productivity estimates is that, if we do not account in their estimation for the impact of different factors of production on the firm's total sales, they may just become an imperfect proxy of these total sales, which would cause our estimate of the elasticity of exports with respect to demand-driven changes in domestic sales to be biased downwards. We should however point out that our measures of productivity are far from being perfectly correlated with the firm's total sales; specifically, this correlation is 0.56 for our baseline approach and 0.22 for our alternative approach. The higher correlation of our baseline approach is consistent with it partly accounting for the firm's usage of material inputs.

Perhaps reflecting the lower correlation between our alternative productivity proxy and the firm's total sales, the OLS estimator in column 3 reveals a *positive* partial correlation between exports and domestic sales. However, the IV elasticity in column 4 is again negative and, though it is significantly lower in absolute value than in our baseline specification (see column 2), it still implies a sizeable substitution between domestic sales and exports at the firm level.

A second concern with our productivity estimates is that, because we do not observe separately prices and quantities for each firm, they may capture not only the firm's actual productivity but also the firm's demand shifter. Specifically, this would be a concern if our productivity estimates were implicitly already controlling for the impact of our instrument. There is however no empirical

[36] In Table G.11, we additionally control for the change in the number of bank offices per capita in the municipality of location of a firm and for the change in firm-level short-term liabilities over total liabilities. We interpret the first of these two variables as an alternative proxy for firms' financial constraints and the second one as a way of partly capturing the potential role of international trade credit in facilitating the growth of exports in municipalities that were hard-hit by the financial crisis. In Table G.11, we also control for the change in the value of land (measured at the municipality level). The estimate of our key parameter is robust to the inclusion of these additional controls.

Table 9: Alternative TFP Measures

Dependent Variable:	ΔLn(Exports)			
TFP Measure:	*TFP Sales*		*TFP Value Added*	
	OLS	2SLS	OLS	2SLS
	(1)	(2)	(3)	(4)
ΔLn(Domestic Sales)	-0.292^a	-1.607^a	0.020	-1.057^a
	(0.032)	(0.248)	(0.038)	(0.197)
ΔLn(Average Wages)	-0.723^a	-1.387^a	-0.753^a	-1.024^a
	(0.072)	(0.151)	(0.070)	(0.100)
ΔLn(TFP Sales): Baseline	1.535^a	2.810^a		
	(0.057)	(0.213)		
ΔLn(TFP Value Added)			1.014^a	1.338^a
			(0.076)	(0.096)
Observations	8,009	8,009	8,009	8,009
F-Statistic		122.44		66.29

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.

evidence of this being relevant in our data: the correlation between our gravity-based instrument for location-specific demand shocks and log changes in our productivity measures is actually negative (it is -0.18 for our baseline productivity estimate and -0.33 for our alternative one).

6.5 Additional Robustness Tests

We finally succinctly discuss a number of additional robustness tests. To save space, these results are reported in Appendices G and H, and we focus throughout on the gravity-based instrument. First, we study in columns 1 to 3 of Table G.5 how our results are affected when restricting our estimation sample according to different criteria. First, we exclude Spanish subsidiaries of foreign multinationals, since both their exports and domestic sales may react to local demand shocks differently than for other firms in Spain. Second, we restrict our attention either to firms with a single manufacturing establishment or even to firms with a single establishment, as multi-establishment firms might have production plants in locations other than the one where they are incorporated and, consequently, may react to local demand shocks in the headquarters' location very differently from single-establishment firms. No matter which of these sample restrictions we implement, the results are not significantly affected.[37] In columns 4 and 5 of Table G.5, we also verify that our results are not materially affected when defining the bust period as 2010-2013 or 2011-13, instead of 2009-13.[38]

[37]Relatedly, we have used information on the location of all car assembly plants in Spain (many of which are part of multi-plant firms), and have confirmed that our results are robust to excluding all provinces where these plants are located.

[38]The motivation for exploring these alternative definitions of the bust period is that the acceleration in the growth rate of Spanish exports starts in 2010 (see Figure 2) but, relative to other countries in the euro area, this acceleration

In Tables G.6, G.8, and G.10, we modify our baseline regression specification so as to: (i) include province or province-sector fixed effects; (ii) cluster standard errors at various levels other than province, and (iii) weight the observations according to different criteria (number of years exporting, log average sales during the boom period, log average employment during the boom period, and log average assets during the boom period). The inclusion of province and province-sector fixed effects and the weighting of observations have a very small effect on our estimates. Some forms of clustering (particularly two-way clustering by province and sector) tend to yield larger our standard errors, but our key estimates remain significant at the 1% level.

Next, in Tables G.11 and G.12, we report specifications that control for a variety of additional firm- or municipality-level supply factors, and specifications that control for firm-specific log average sales, log average employment, log average assets or average export-to-sales ratio during the boom period. When introducing these controls, the estimates of the elasticity of interest are never lower than -1.4 or higher than -1.9 in absolute value. In Tables G.18 and G.19, we reproduce our findings when aggregating the firm-level data at the municipality-sector level. When estimating our baseline specification at the municipality-sector level, we obtain lower (in absolute value) elasticities of exports with respect to domestic sales, but the qualitative nature of our findings remains unaffected.

In Appendix H, we explore alternative strategies to identify the potential substitutability of domestic and export markets. First, in Appendix H.1, we follow the identification approach implemented Berman et al. (2015) to estimate the causal impact of demand-driven changes in exports on domestic sales, swapping then the role that these two variables play in our main specification. Due to data restrictions (see section 3.3), we can only carry out such analysis for the period 2002-07. Consistently with our main findings, and contrary to the results in Berman et al. (2015) using French data, we find a *negative* effect of demand-driven changes in exports on domestic sales. Second, in Appendix H.2, we re-estimate the parameters of our main regression specification but explore alternative instrumentation strategies that focus on the deep roots of the differential fall in demand across Spanish regions. More specifically, we posit that, relative to the boom years, municipality-level demand shocks were larger (i) in municipalities with lower housing supply elasticities (in which house prices grew disproportionately during the boom years), (ii) in municipalities with a larger pre-crisis contribution of the construction sector to total labor income, and (iii) in provinces that experienced larger declines in tourism during the bust years.[39] We then weigh the municipality-specific demand shocks in (i) and (ii) following the same procedure as in our baseline gravity-based instrument. The second-stage estimates of our parameter of interest computed using these alternative instruments are all negative and generally a bit lower in absolute value than our baseline one; however, one should interpret these estimates with caution, as these instruments are generally not as strong as our baseline one.[40]

starts in 2011 (see Figure 1).

[39]The construction and tourist sectors are among the ones that experienced the largest reduction in total sales and employment in the bust relative to the boom. Regions more exposed to these sectors are likely to have experienced a larger drop in demand for manufactured goods.

[40]When adding each of these instruments one by one to our baseline gravity-based instrument, standard tests of overidentifying restrictions fail to reject at typically used significance levels the null hypothesis that our instruments are jointly valid.

7 Model With Increasing Marginal Costs

Our empirical results suggesting a negative impact of demand-driven changes in domestic sales on changes in exports are in contradiction with the framework described in section 2. In this section, we show how a simple extension of that framework can rationalize our empirical results.

7.1 Model With Increasing Marginal Costs: Estimating Equation

The theoretical environment we consider here is identical to that in section 2, except that the cost structure in equation (2) is now replaced with a total variable cost of producing Q_{id} units of output for the domestic market and Q_{ix} units of output for the foreign market given by

$$\frac{1}{\widetilde{\varphi}_i}\omega_i \frac{1}{\lambda+1}\left(\tau_d Q_{id} + \tau_x Q_{ix}\right)^{\lambda+1}, \quad \lambda \geq 0, \tag{14}$$

where $\tau_d Q_{id} + \tau_x Q_{ix}$ denotes firm i's total output in the presence of iceberg trade costs in the domestic (τ_d) and foreign (τ_x) markets. Notice that the parameter λ governs how steeply marginal costs increase with output. When $\lambda = 0$, marginal costs are constant and equation (14) reduces to our previous expression in equation (2). We show in Appendix A that the cost function in equation (14) can be derived in a model in which the firm's production function is a Cobb-Douglas aggregator of a fixed or pre-determined input and a flexible and static input; without loss of generality, we can refer to these two inputs as capital and labor, respectively. Under this microfoundation, the parameter λ is decreasing in the elasticity of output with respect to the flexible factor, and λ is equal to 0 when this elasticity is equal to one. Note also that we denote firm productivity with the new notation $\widetilde{\varphi}_i$ (rather than φ_i), as the microfoundation in Appendix A shows that this productivity level $\widetilde{\varphi}_i$ depends not only on the TFP level φ_i but also on the stock of fixed factors.

Solving for the optimal level of exports by firm i under the cost function in equation (14), and taking log differences, we obtain

$$\begin{aligned}\Delta \ln R_{ix} \quad = \quad & (\sigma-1)\left[\Delta \ln \xi_{ix} + \Delta \ln \widetilde{\varphi}_i - \Delta \ln \omega_i\right] - (\sigma-1)\left(\Delta \ln \tau_{sx} - \Delta \ln P_{sx}\right) + \Delta \ln E_{sx} \\ & - (\sigma-1)\,\lambda \Delta \ln \left(\tau_d Q_{id} + \tau_x Q_{ix}\right),\end{aligned} \tag{15}$$

which is analogous to equation (4) except for the last term, which reflects the effect of total output on marginal production costs.[41] Next, note that, due to constant mark-up pricing, we can write

$$\ln\left(\tau_d Q_{id} + \tau_x Q_{ix}\right) = \ln\left(\frac{\tau_d R_{id}}{P_{id}} + \frac{\tau_x R_{ix}}{P_{ix}}\right) = \ln\left(R_{id} + R_{ix}\right) - \ln\left(\frac{\sigma\omega_i\left(\tau_d Q_{id} + \tau_x Q_{ix}\right)^{\lambda}}{(\sigma-1)\widetilde{\varphi}_i}\right). \tag{16}$$

Solving for $\ln\left(\tau_d Q_{id} + \tau_x Q_{ix}\right)$, plugging this expression into equation (15), and imposing the same

[41] All expressions in this section implicitly assume there are only two markets, one domestic market and one foreign market. For an extension of this model to a setting with multiple domestic and foreign markets, see Appendix E.2.

decomposition as in equation (5), we then find that:

$$\Delta \ln R_{ix} = \gamma_{sx} + \frac{(\sigma-1)}{1+\lambda}\delta_{\varphi}\Delta \ln \widetilde{\varphi}_i^* - \frac{(\sigma-1)}{1+\lambda}\delta_{\omega}\Delta \ln \omega_i^* - \frac{(\sigma-1)\,\lambda}{1+\lambda}\Delta \ln (R_{id}+R_{ix}) + \varepsilon_{ix}, \quad (17)$$

where $\varepsilon_{ix} \equiv u_{ix}^{\xi} + ((\sigma-1)/(1+\lambda))(u_i^{\varphi} - u_i^{\omega})$. This equation is analogous to equation (10), except that it features the log difference of *total* sales (instead of domestic sales) on the right-hand side, and that it calls for the inclusion of the firm's stock of fixed factors as an additional control. The intuition for the need to include the change in total sales rather than in domestic sales as an explanatory variable is straightforward: marginal costs of production are increasing in total output, not just output destined for the domestic market.

Estimating equation (17) via OLS is problematic not just for the reasons identified in section 2, but also because the fact that the log change in total sales naturally depends on the log change in exports implies that any unobserved determinant of exports accounted for by the regression residual will be correlated with our covariate of interest and, thus, will bias the OLS estimate of $(\sigma-1)\,\lambda/\,(1+\lambda)$. Importantly, because the regression residual in equation (17) depends on the same terms as that in equation (10) (i.e., unobserved productivity, factor costs and export demand shifters), a TSLS estimator based on our instrument will deliver consistent estimates of this regression coefficient as long as the identification assumptions outlined in section 2.1 hold. Consequently, the threats to the validity of our instrument discussed in section 4.2 also apply here.

In Table 10, we present OLS and TSLS estimates of the regression coefficients in equation (17). In columns 1 to 3 of panels A and B, we include in the regression specification the same controls as in Tables 2 and 3. In columns 4 and 5, we additionally control for the change in the stock of capital, as indicated by the micro-foundation in Appendix A. As expected, the OLS estimates in column 1 indicate a strong positive correlation between exports and total sales, even when controlling for sector fixed effects and for our measures of firms' average wages and TFP. The first-stage results in column 2 indicate that both the local and the gravity-based instruments are strong predictors of a firm's total sales, with an F-stat of 28.99 and 75.00, respectively. The second-stage elasticities of exports to total sales in column 3 are negative and significant and stand at a value of -2.038 and -2.374, respectively. Adding the boom-to-bust log change in the firm's stock of physical capital does not affect significantly the first-stage nor the second-stage results. Thus, henceforth, we treat the estimates in column 3 – specifically the one in panel B – as our baseline estimates.

To understand the magnitude of our estimates, take a firm with an initial export share of 16.2% (which corresponds to the median export share during the boom in our sample of 8,009 continuing exporters). Suppose that, due to a drop in demand, this firm experiences a 1% drop in its domestic sales. Our estimated elasticity of exports to domestic sales in Table 3 indicates that, other things equal, the firm should see its exports increase by 1.6%. This also implies that the firm's total sales will decrease by $83.8\% \times 1\% + 16.2\% \times (-1.6\%) = 0.58\%$. For this change in total sales, our estimated elasticity of exports to total sales in panel B of Table 10 suggests an implied increase in exports of $0.58\% \times 2.374 = 1.4\%$, which is close to the 1.6% increase predicted by the estimates in Table 3. This demonstrates that our IV results in Tables 3 and 10 deliver congruent estimates for

Table 10: Intensive Margin with Total Sales

	Panel A: Local Instrument				
Dependent Variable:	ΔLn(Exp)	ΔLn(TotSales)	ΔLn(Exp)	ΔLn(TotSales)	ΔLn(Exp)
	(1)	(2)	(3)	(4)	(5)
	OLS	1st Stage	2nd Stage	1st Stage	2nd Stage
ΔLn(Total Sales)	0.734^a		-2.038^a		-2.253^a
	(0.037)		(0.633)		(0.687)
ΔLn(TFP)	0.494^a	1.033^a	3.339^a	0.987^a	3.404^a
	(0.058)	(0.025)	(0.639)	(0.025)	(0.662)
ΔLn(Average Wages)	-0.205^a	-0.499^a	-1.579^a	-0.483^a	-1.633^a
	(0.054)	(0.032)	(0.312)	(0.031)	(0.330)
Δ Ln(Vehicles p.c. in municipality)		0.272^a		0.255^a	
		(0.051)		(0.047)	
Δ Ln(Stock of Capital)				0.102^a	0.348^a
				(0.009)	(0.076)
Observations	8,009	8,009	8,009	8,009	8,009
Sector FE	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes
F-statistic		28.99		29.27	
	Panel B: Gravity-based Instrument				
Dependent Variable:	ΔLn(Exp)	ΔLn(TotSales)	ΔLn(Exp)	ΔLn(TotSales)	ΔLn(Exp)
	(1)	(2)	(3)	(4)	(5)
	OLS	1st Stage	2nd Stage	1st Stage	2nd Stage
ΔLn(Total Sales)	0.724^a		-2.374^a		-2.590^a
	(0.050)		(0.526)		(0.606)
ΔLn(TFP)	0.509^a	1.063^a	3.690^a	1.015^a	3.739^a
	(0.055)	(0.026)	(0.482)	(0.027)	(0.539)
ΔLn(Average Wages)	-0.217^a	-0.509^a	-1.750^a	-0.493^a	-1.801^a
	(0.063)	(0.043)	(0.250)	(0.041)	(0.282)
ΔLn(Dist-Pop-Weighted Vehicles p.c.)		0.888^a		0.838^a	
		(0.103)		(0.107)	
ΔLn(Stock of Capital)				0.101^a	0.382^a
				(0.009)	(0.067)
Observations	8,009	8,009	8,009	8,009	8,009
Sector FE	Yes	Yes	Yes	Yes	Yes
F-statistic		75.00		61.30	

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.

the response of exports to local demand shocks.[42]

With an estimate of the demand elasticity σ in hand, it is easy to infer an estimated value of

[42] In Appendix I, we present results for specifications analogous to those in Tables 2 to 9, with the only difference being that the boom-to-bust log change in *total* rather than *domestic* sales is included as right-hand-side variable The conclusions discussed in section 6 are generally corroborated by the results reported in Appendix I.

λ from the estimates in Table 10. Specifically, given the estimates in column 3 of panel B, we can compute an estimate of λ as $\hat{\lambda} = 2.374/(\sigma - 1 - 2.374)$. For $\sigma = 6$, we obtain $\hat{\lambda} = 0.90$, which indicates a significant departure from constant marginal costs.

To complement these findings, in Appendix I we conduct additional tests of several implications of the vent-for-surplus mechanism. These tests are based on the idea that one should expect the increase in exports in reaction to a common demand-driven drop in domestic sales to be 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. The results in Appendix Table I.8 confirm that the elasticity of the change in exports to changes in total sales is higher for firms having lower output elasticities with respect flexible inputs, although few of the results are statistically significant at standard levels.

8 Quantification

In this final section, we use the extended model with increasing marginal costs to evaluate the quantitative importance of the "vent-for-surplus" channel for explaining the remarkable growth in Spanish exports during the period 2009-13.

To achieve this goal, we implement a three-step procedure. First, we use the model described in section 7 to trace the impact that arbitrary counterfactual changes in Spain's sectoral domestic aggregate demand shifters – or $B_{sd} \equiv E_{sd}/P_{sd}$, where E_{sd} denotes the total expenditure in Spain in sector s and P_{sd} is the price index of sector s in Spain – would have had on Spanish aggregate exports, domestic and total sales. In doing so, we quantify the impact that these demand changes had on firms' sales exclusively via the "vent-for-surplus" channel (i.e., the movement of firms along their marginal cost curves). Second, we perform a variance decomposition of the observed change in firms' total sales with the aim of informing the precise extent to which B_{sd} fell during the crisis. With this number at hand, in our last step, we compute our model's prediction for the counterfactual change in aggregate exports that would have been observed if there had been no change in aggregate demand shifters between the boom and the bust. If this number is small relative to the observed boom-to-bust growth in exports, we can conclude that the "vent-for-surplus" channel must have been an important determinant of the Spanish export miracle during the period 2009-13.

We next explain these steps in more detail, and also highlight the underlying assumptions behind them.

Step 1: Computing counterfactual changes in aggregate domestic sales and exports for given changes in the sectoral demand shifters $\{B_{sd}\}_{s=1}^{S}$. Through the lens of our model, we capture the domestic demand changes that affected Spanish firms between the boom and the bust as changes in the sectoral demand shifters $\{B_{sd}\}_{s=1}^{S}$. According to our model, changes in B_{sd} for a sector s determine changes in the residual demand function that each firm in s faces and, thus, from the perspective of each individual firm, are purely demand shifters.[43]

[43] In models that endogenize these demand shifters, changes in B_{sd} could be due to a variety of factors. For example, in a model in which consumers have Cobb-Douglas preferences over the S manufacturing sectors we consider in our

The system of equations that allows us to map relative counterfactual boom-to-bust changes in sectoral demand shifters to changes in aggregate domestic sales and exports (and, thus, total sales) boils down to three sets of equations. The first one maps the boom-to-bust counterfactual change in exports of each firm to its counterfactual change in domestic sales. The second one maps the boom-to-bust counterfactual change in domestic sales of each firm to its counterfactual change in exports and to the counterfactual changes in the aggregate demand shifter and price index in the firm's sector. Finally, the third one maps the counterfactual change in the price index of each sector to the counterfactual change in the aggregate demand shifter of the corresponding sector. We describe these three sets of equations in detail in Appendix J.1.

Using these equations, we compute the impact that relative counterfactual changes in sectoral demand shifters – $(B'_{sd1}/B_{sd0})(B_{sd1}/B_{sd0})^{-1}$ – have on firm-specific counterfactual changes in exports and domestic sales. We then aggregate these firm-specific counterfactual changes across all firms active in the boom and bust periods and thus construct counterfactual changes in aggregate domestic sales and exports. When performing our baseline quantification, we set $\sigma = 5$, which is a central value in the range of estimates used in the international trade literature (see Head and Mayer, 2014), and $(\sigma - 1)\lambda/(1 + \lambda) = 2.374$, which corresponds to the estimate reported in column 3 of Table 10.[44]

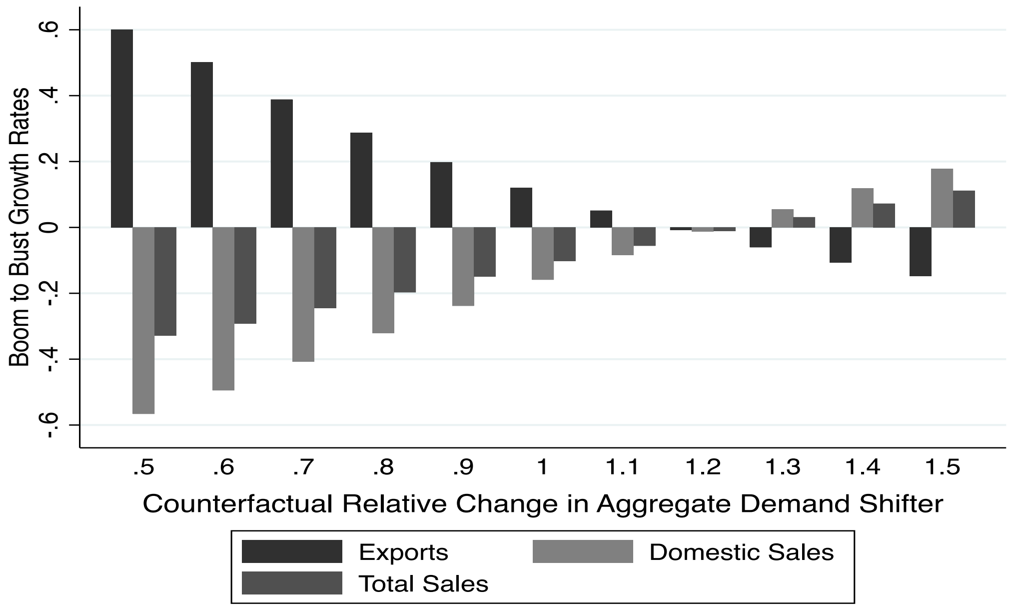
Among all the relative counterfactual changes in sectoral demand shifters we could consider, we focus on relative changes that are constant across sectors; i.e., (B'_{sd1}/B_{sd0}) $(B_{sd1}/B_{sd0})^{-1}$ equals a constant, which we denote as Γ_B. In Figure 4, we plot the changes in aggregate exports, domestic sales, and total sales predicted by our model when Γ_B takes different values between 0.5 and 1.5. For the set of firms that we use in our counterfactual analysis, we observe in our data that, between the boom and the bust, aggregate domestic sales dropped 15.91%, exports grew by 11.99%, and total sales dropped by 10.23%. These are the values that our counterfactual analysis naturally generates when we set the relative counterfactual change in the aggregate demand shifter of every sector s to equal 1; i.e., $\Gamma_B = 1$. If the value of the aggregate demand shifters in the bust had been 50% smaller than they actually were (i.e., $\Gamma_B = 0.5$), our model predicts that aggregate domestic sales would have dropped by 56.64% and aggregate exports would have increased by 60.1%. In this case, aggregate total sales would have dropped by 32.87%. Conversely, if it had been 50% larger (i.e., $\Gamma_B = 1.5$), aggregate domestic sales would have grown by 17.82% and aggregate exports would have dropped by 14.76%, and the result would have been an 11.18% growth in total sales.

The above predictions rely on several key assumptions. First, when computing the aggregate change in domestic sales and exports that would have been observed for different changes in the demand shifter, we effectively maintain the boom-to-bust changes in the supply parameters $(\varphi_i, \omega_i, \tau_{sx}, \tau_{sd})$ and in the firms' idiosyncratic demand shifters (ξ_{id}, ξ_{ix}) at their realized values, independently of the size of the counterfactual demand change. Admittedly, the simultaneous move

analysis and other 'outside' sectors, decreases in the share of spending on manufacturing goods (e.g., a decrease in the Cobb-Douglas parameters associated with manufacturing sectors) would reduce the demand shifters $\{B_{sd}\}_{s=1}^{S}$. In a model in which consumers own houses, demand shifters would respond to wealth effects associated to changes in housing prices, which may themselves reflect changes in expectations of future housing demand and supply shocks.

[44] In Appendix J.4, we explore how robust our results are to alternative values of $(\sigma - 1)\lambda/(1 + \lambda)$.

Figure 4: Impact of Aggregate Demand Shocks



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 computed using equations (J.4), (J.5) and (J.11), described in Appendix J.1. Given these counterfactual growth rates in export and domestic sales, we compute the counterfactual growth rate in total sales as $(R'_{ix1}/R_{ix0})\chi_{i0}+(R'_{id1}/R_{id0})(1-\chi_{i0})$.

of all Spanish firms along their marginal cost curves could in principle have impacted equilibrium wages and, more generally, equilibrium input prices. Although we do not take these effects into account in our baseline quantification, as a robustness check, we present in Appendix J.5 results in which we illustrate how our main predictions are affected if we allow firms' wages to change as firms move along their marginal cost curves.[45]

Second, our counterfactual calculations rely on the assumption that, in every sector, Spain is a small open economy; thus, counterfactual boom-to-bust changes in Spanish aggregate demand affect neither the boom-to-bust change in the foreign price index nor the boom-to-bust change in foreign firms' marginal production costs.[46]

Third, consistently with the estimates presented in Table 5, we assume that firms do not change their export status in reaction to changes in the domestic demand shifters. Consequently, we focus on quantifying the impact of counterfactual changes in aggregate demand on the aggregate volume of exports of those firms that export during both the boom and the bust periods.

Step 2: Decomposing the variance of total sales. Our ultimate goal is not to indicate how aggregate

[45]The movement of each firm along its marginal cost curve will impact its output price and, consequently, the simultaneous move of all firms along their marginal cost curves will impact the domestic sectoral price index P_{sd} for every sector s. This effect *is* taken into account in our quantification.

[46]In a model with increasing marginal costs such as ours, the invariability of the foreign price index to changes in Spanish aggregate demand shifters implies that Spain is a small exporter to the rest of the world, and the invariability of foreign firms' marginal production costs implies that Spain is a small importer from the rest of the world.

exports, domestic and total sales react to arbitrary counterfactual domestic demand changes, but to predict the change in these variables that we would have observed if demand shifters in Spain had remained constant between the boom and the bust. Doing so requires measuring the extent to which Spanish domestic demand actually fell between these two periods. Given the lack of firm-specific output prices in our dataset, our measure of Spain's boom-to-bust drop in domestic demand shifters uses as key input a decomposition of the variance of firms' boom-to-bust changes in total sales into changes due to demand factors and changes explained by supply factors. This variance decomposition exercise is analogous to that implemented in Autor et al. (2013) with the purpose of measuring the impact that productivity growth in China had on the growth in US imports from China during the period 1990-2007. Specifically, we use equation (17) to decompose the variance of $\Delta \ln(R_i)$, with $R_i \equiv R_{id} + R_{ix}$, into a component due to firms' marginal cost and export demand shifters, and a component attributed to factors orthogonal to these shifters (see Appendix J.2 for details). We find the contribution of the combination of marginal cost and export demand shifters to be 59%, and that of factors orthogonal to it to be 41%. On the basis of this number, we infer that 41% of the 10.23% drop in total sales between the boom and the bust periods was due to domestic demand factors.[47]

Step 3: Quantification results. Given the finding that 41% of the 10.23% boom-to-bust drop in total sales was due to changes in demand, we use the counterfactual results described in *Step 1* to find the value of Γ_B for which our model predicts a drop in total sales of $6.04\% = (1 - 41\%) \times 10.23\%$. Intuitively, this is the drop in total sales that we would have observed if aggregate demand shifters had not changed between the boom and the bust. Our model predicts a drop in total sales of 6.04% if $\Gamma_B = 1.09$.[48] For this value of Γ_B, our model predicts that exports would have grown in 5.79%. 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.[49]

9 Conclusion

In this paper, we provide evidence suggesting that export and domestic sales decisions are interdependent at the firm level. Faced with a severe domestic slump during the Great Recession, Spanish producers appear to have experienced a decline in their short-run marginal production costs, with this gain in competitiveness translating into an increase in their sales in foreign markets. We cir-

[47]Being precise, our decomposition reveals that 41% of the variance of the changes in firms' total sales is due to *any* factor orthogonal to firms' marginal cost shifters and export demand shocks. Thus, our conclusion that changes in demand shifters explain 41% of the variance of the changes in firms' total sales implicitly assumes that these shifters are the only determinants of firms' total sales whose boom-to-bust changes are orthogonal to changes in firms' marginal cost shifters and export demand shocks.

[48]If it were true that, as we impose in our analysis, the boom-to-bust change in the aggregate demand shifters $\{B_{sd}\}_{s=1}^{S}$ was the same across all manufacturing sectors $s = 1, \ldots, S$, then we can infer from our analysis that the aggregate demand shifters fell between the boom and the bust in $1 - 1/\Gamma_B^* = 1 - 1/1.09 = 1 - 0.9174 = 8.26\%$.

[49]Our analysis accounts for the changes in trade costs that took place between the boom and the bust. In Appendix J.3, we explore further the role that trade costs may play in inhibiting firms' ability to exploit the vent-for-surplus mechanism. Our results show that trade costs play an important role: if trade costs in the bust had been 10% larger than they were, then total sales would have dropped by 13.97% (while they dropped by 10.23% in the data).

cumvent the inherent difficulties associated with establishing a causal link between demand-driven changes in domestic sales and exports by exploiting geographic variation in the incidence of the Great Recession in Spain.

Our empirical findings are inconsistent with international trade models featuring constant markups and technologies with constant marginal costs of production. We rationalize and interpret our results through the lens of a model with increasing marginal costs, and show that the "vent-for-surplus" mechanism is powerful enough to explain approximately half of the growth in Spanish exports in the period 2009-13. Although there are a few singular aspects of the Spanish experience in the years around the Great Recession that may make the "vent-for-surplus" mechanism particularly important in this context (e.g., the large boom in investment preceding the bust, or the sclerotic nature of the Spanish labor market), we think that some of the insights and methodology in our paper can be transported to other countries that experienced severe domestic slumps.

Our paper also offers a new perspective on the literature studying interdependencies in the extensive and intensive margins of trade (e.g., Antràs et al. 2017; or Morales et al. 2019). Due to data limitations, we have restricted our analysis to the study of interdependencies between the domestic market and a single (aggregate) export destination, and we have modeled these interdependencies as arising exclusively from an increasing marginal cost function. In such a case, the firm's profit function is submodular in the extensive margin of trade, a feature that would be preserved in a model with multiple export markets, as shown in Appendix E.2. With access to data on firms' exports by destination market, and borrowing tools from Arkolakis and Eckert (2017), one could estimate the key parameters of a multi-country version of our model, and thus explore interdependencies also in the intensive and extensive margin of trade *across export markets*. Even more ambitiously, with better data on prices, one could potentially expand our analysis to explore the extent to which endogenous markup adjustments (see De Loecker et al., 2016) or price stickiness in both the domestic and export prices (Gopinath and Rigobon, 2008; Nakamura and Steinsson, 2008, 2013) affect the way in which a firm's exports react to domestic demand shocks. We leave the study of these questions for future research.

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A Convexity of the Short-run Marginal Cost Function

Suppose a firm's production function depends on fixed or pre-determined input K_i and a flexible and static input L_i. Let us refer to the former as capital and the latter as labor. Assuming a Cobb-Douglas technology in capital and labor, the cost minimization problem of a firm with productivity φ_i seeking to produce a total amount of output Q_i can be expressed as:

$$\begin{aligned} & \min \quad \omega_i L_i \\ s.t. \quad & \varphi_i K_i^{\alpha_K} L_i^{\alpha_L} \geq Q_i, \end{aligned}$$

where ω_i denotes the nominal wage that firm i faces, and α_K and α_L denote then the output elasticities with respect to capital and labor, respectively. The first-order condition of the cost-minimization problem of the firm delivers

$$\begin{aligned} \omega_i &= \mu \alpha_L \frac{Q_i}{L_i} \\ \varphi_i K_i^{\alpha_K} L_i^{\alpha_L} &= Q_i, \end{aligned}$$

where μ denotes the Lagrange multiplier on the constraint $\varphi_i K_i^{\alpha_K} L_i^{\alpha_L} = Q_i$. After solving for L_i in the second of these equalities, we can rewrite the short-run total costs as a function of output, Q_i, as follows

$$\omega_i L_i = \omega_i \left(\varphi_i K_i^{\alpha_K}\right)^{-\frac{1}{\alpha_L}} (Q_i)^{\frac{1}{\alpha_L}} .$$

Note that, unless $\alpha_L = 1$, this short-run total cost function is not linear in the total output Q_i. More specifically, as long as $0 < \alpha_L < 1$, this short-run cost function will be convex in Q_i. Using $\widetilde{\varphi}_i$ to denote a shifter of the short-run costs, and using λ to denote the deviation of the output elasticity of the short-run cost function relative to the case in which this function is linear in Q_i, i.e.,

$$\begin{aligned} \widetilde{\varphi}_i &= \alpha_L \left(\varphi_i K_i^{\alpha_K}\right)^{\frac{1}{\alpha_L}} \\ \lambda &= \frac{1-\alpha_L}{\alpha_L}, \end{aligned}$$

we can rewrite the short-run total costs as

$$\omega_i L_i = \frac{1}{\widetilde{\varphi}_i} \omega_i \frac{1}{1+\lambda} (Q_{it})^{1+\lambda}.$$

The elasticity of the short-run total costs with respect to output is thus

$$\frac{\partial \ln(\omega_i L_i)}{\partial \ln(Q_{it})} = 1 + \lambda.$$

Note that, the lower the value of α_L (i.e., the lower the elasticity of output with respect to the flexible input), the larger the elasticity with respect to output of the short-run total cost function. The curvature of the total cost schedule is thus crucially shaped by the parameter determining the elasticity of output with respect to the flexible input.

B Data Appendix

B.1 Macroeconomic Data

Data on Spanish unemployment, real GDP, internal demand, private final consumption expenditure and exports of goods come from the Spanish National Statistical Office (*Instituto Nacional de Estadística*). Data on merchandise exports and real GDP shares for the countries that belong to the European Monetary Union come from the AMECO Dataset (i.e., annual macro-economic database of the European Commission's Directorate General for Economic and Financial Affairs). Data on unit labor costs in the manufacturing sector for Spain and the European Monetary Union were obtained from the Bank of Spain (*Banco de España*), the Eurosystem, and the OECD dataset on Productivity and ULC by main economic activity. We use the input-output tables produced by the Spanish National Statistical Office for the year 2005 to identify the interlinkages across industries (e.g., the two top leading input providers or two top leading buying industries of the vehicles manufacturing industry discarded in the robustness analysis described in Table 6).

B.2 Construction of the Commercial Registry Dataset

As described in section 3.3, our main source of firm-level data is the Commercial Registry (*Registro Mercantil Central*), which contains annual financial statements of around 85% of registered firms in the non-financial economy. We collect data from two separate sources to construct our own firm-level dataset: (i) the *Central de Balances* dataset from the Bank of Spain, and (ii) *Sabi*, from Informa, a private company. The Bank of Spain made an effort to expand and treat the information for small firms gathered in the Commercial Registry but the Central de Balances dataset does not cover the universe of private-sector firms. In particular, this dataset excludes, mainly, medium and large firms that submit information after the regular submission deadline or that do not use a digital support. Conversely, the Informa dataset puts special emphasis on compiling information on large and medium-sized firms that submit their statements either late or on paper. We combine the information in these two datasets to take advantage of their complementarities in order to maximize the coverage of the resulting database. A detailed description of how we combine the two sources to construct our firm-level dataset can be found in Almunia et al. (2018).

In terms of the sectoral disaggregation of the data, note that NACE (*Nomenclature générale des activités économiques dans les Communautés Européennes*) is the European statistical classification of economic activities. It classifies manufacturing firms into 24 different sectors. Some firms move to a different zip code or change their sectoral classification during the period of analysis. In the boom-to-bust regressions reported later, we assign to these firms a fixed zip code and sector using their most frequent value in each case. A firm's zip code corresponds to the location of its headquarters.

B.3 Foreign Transactions Dataset

As described in section 3.3, until 2014 the Bank of Spain required all financial institutions and a set of large companies to report all foreign transactions, including imports, exports and other financial transactions. Until 2007, there is information for each transaction on the country of destination (or origin). However, from 2008 to 2013, the Bank of Spain relaxed this requirement and allowed reporting institutions to group multiple transactions into a single reported transaction. In those cases, the country of destination (or origin) reflected in the data corresponds to the country of the largest transaction in that group. Similarly, the product code reported corresponds to the largest transaction as well. This implies that one cannot analyze changes in exports or imports by country of destination (or origin) nor by product in a consistent way for periods spanning around 2008. The foreign transactions registry collected by the Bank of Spain was discontinued in early 2014. Since then, the Bank of Spain's monitoring of foreign transactions mainly relies on aggregate data built from transaction-level information that is provided by the Spanish tax administration.

B.3.1 Minimum Reporting Threshold

Between 2001 and 2007, all foreign transactions of more than €12,500 had to be reported to the Bank of Spain. In order to reduce the compliance costs for reporting institutions, the minimum reporting threshold was updated in 2008 to €50,000. From that year onwards, a firm appears in the dataset if it has at least one transaction larger than €50,000 in that year. In order to create a homogeneous sample for the period 2002-2013, we apply the post-2008 minimum reporting threshold to the data from 2002 to 2007, meaning that we only record a positive export flow in a given year for firms that have at least one transaction exceeding €50,000 in that year. This adjustment reduces substantially the number of exporting firms that appear in the data, but the impact on the aggregate amount exported is small.

B.4 Vehicles per Capita and Tax Records of Firm-Level Sales within Spain

The information on the stock of vehicles by both municipality and province is provided by the Spanish Registry of Motor Vehicles. According to Spanish Law, vehicles have to be registered in the municipality where the owner has her permanent residence. This residence status should match the one reported in the municipal census (*Padrón*). In the case of legal entities (business or institutions), vehicles must be registered in the municipality where they undertake their main activity, which should match the location reported to the tax authorities. The census of vehicles at the municipal level is maintained by the General Directorate for Traffic (*Dirección General de Tráfico*). Each city council has the capacity to levy a small fee on the registered vehicles in its municipality (*Impuesto sobre Vehículos de Tracción Mecánica*). This fee usually depends on several criteria such as vehicle power, type of vehicle, pollution level, etc. In aggregate terms, these fees collectively raised tax revenue equivalent to around 0.2% of GDP in 2016.

The information on population both at the municipality and province levels is provided by the Spanish National Statistical Office.

Regarding our data on firm-level sales within Spain, Spanish Tax Law obligates all firms (legal entities) and professionals (natural persons) that undertake economic activities to report detailed information on the transactions with their trading partners. In 2006, this information is collected in Form 347, officially called "Annual information return on transactions with third parties". We work with data for the year 2006 because it is the first year for which a precise and consistent comprehensive digitization of the data is available. In particular, each business must report the monetary value of its individual sales to each trading partner. The reported transactions include all domestic sales to businesses, households and the public sector. The law uniquely exempts mandatory reporting of individual transactions when the annual aggregate sales to a trading partner do not exceed €3,005.06. This tax record of sales to third-parties is a fundamental tool of tax enforcement for both the VAT and the corporate income tax given that transactions included in Form 347 must be reported consistently in both tax returns.

The Spanish Tax Agency (*Agencia Estatal de Administración Tributaria*, AEAT) shared with the Bank of Spain aggregate data on municipality-to-municipality flows for firms in the manufacturing sector, excluding sales of businesses in the auto industry, for the year 2006. In particular, for each municipality where manufacturing sellers are located, the Spanish Tax Agency computed the total amount of sales to any municipality where purchases are made. We thus have access to a matrix of bilateral flows of manufactured goods between Spanish municipalities. This matrix contains data on 485,565 municipality-to-municipality flows, with 2,305 municipalities of origin and 6,623 municipalities of destination. After restricting the sample of municipalities to those observed in our subsample of continuing exporters, the dataset contains a matrix of 412,500 bilateral municipality flows, with 1,224 municipalities of origin and 6,587 municipalities of destination.

Apart from the matrix of aggregate municipality-to-municipality flows, the Spanish Tax Agency also provided us, for each firm in our sample of 8,009 continuing exporters, the 2006 share of its total domestic sales going to each Spanish municipality.

Venting Out: Exports during a Domestic Slump

Miguel Almunia, Pol Antràs, David Lopez-Rodriguez and Eduardo Morales

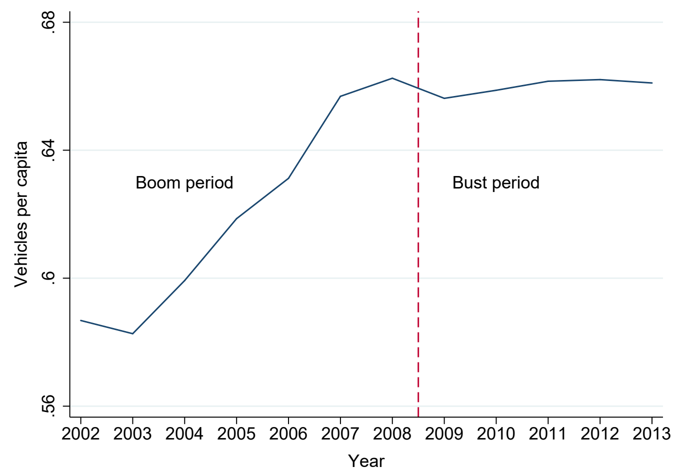
Online Appendix (Not for Publication)

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.

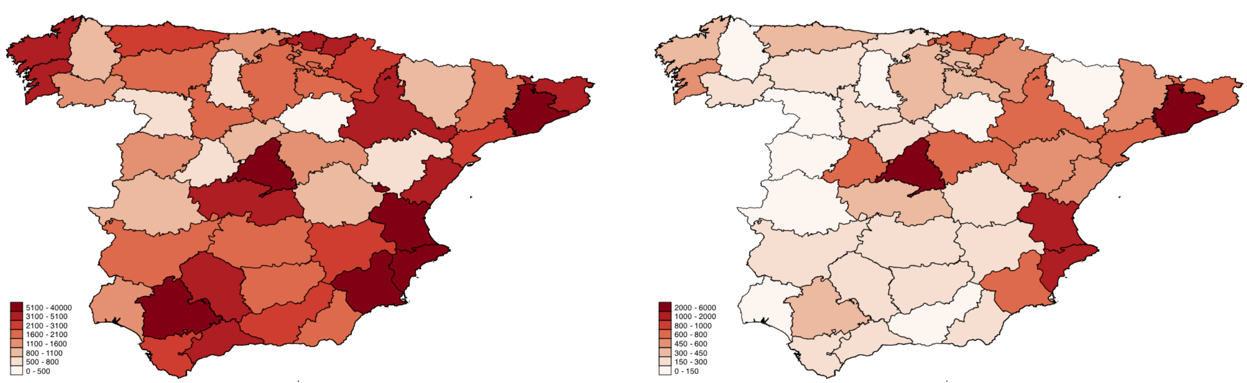
Figure C.1: Stock of Vehicles per Capita in Spain



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.

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



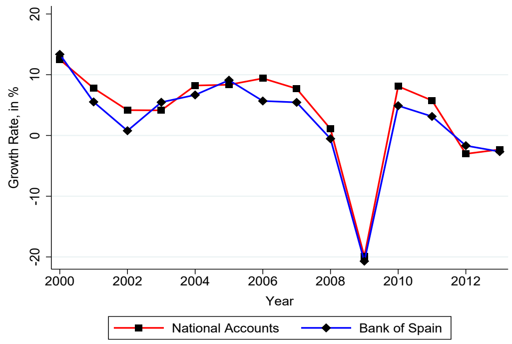
(a) Number of Firms

(b) Number of Exporting Firms

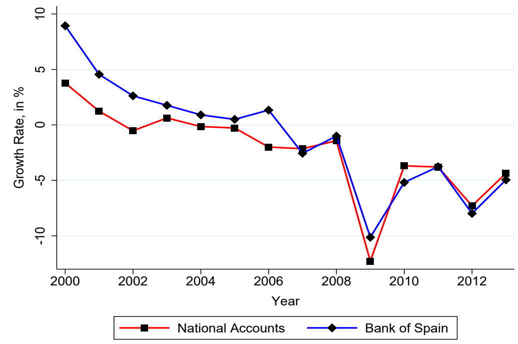
C.3 Data Coverage of Macroeconomic Dynamics

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 and Customs. The figures show that our dataset tracks well the aggregate evolution over time of output, employment, total payments to labor, and exports.

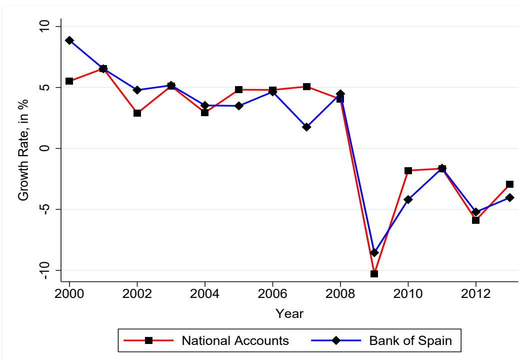
Figure C.3: Output, Employment, Wage Bill and Export Dynamics



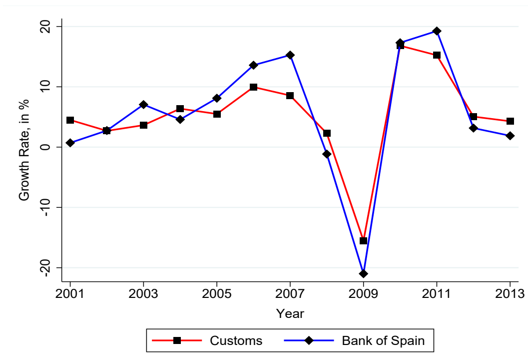
Panel (a): Output



Panel (b): Employment



Panel (c): Wage Bill



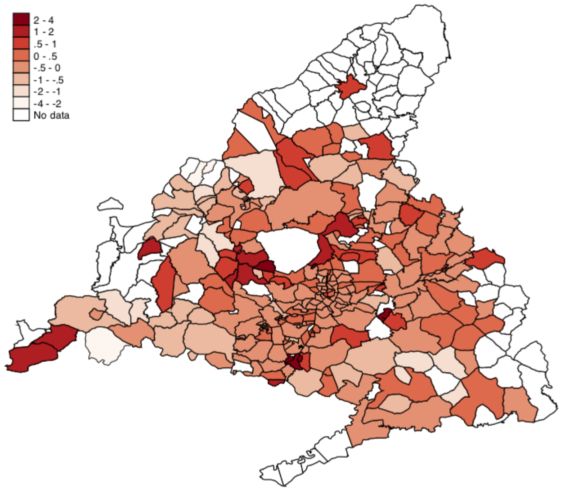
Panel (d): Exports

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

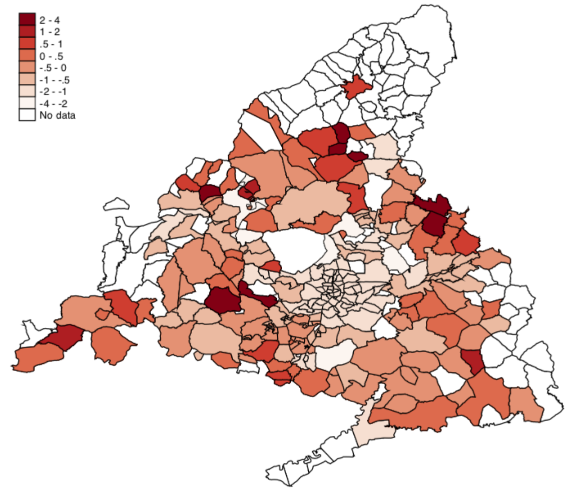
Figure C.4 illustrates variation across zip codes in both the boom-to-bust changes in average manufacturing firm-level domestic sales and the boom-to-bust changes in the number of vehicles per capita. We do so for the case of the two most populated provinces in Spain: Madrid and Barcelona. To facilitate 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.

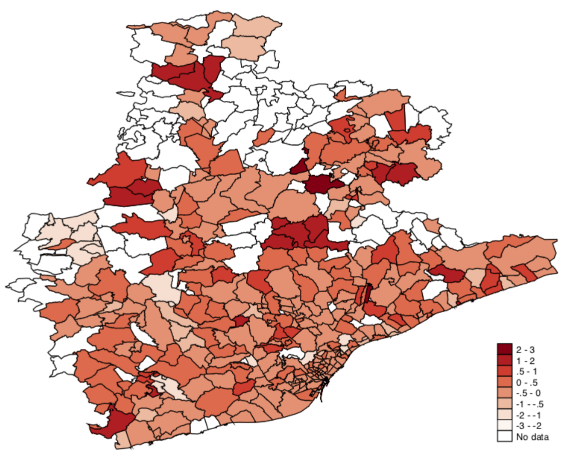
Figure C.4: The Great Recession in Madrid and Barcelona: Variation Across Zip Codes



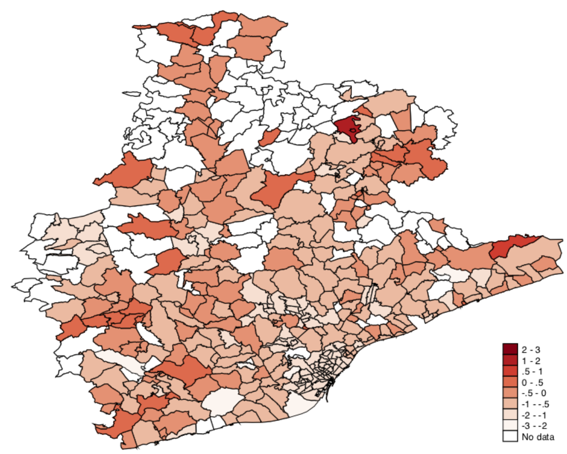
(a) Relative Change in Domestic Sales (Madrid)



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



(c) Relative Change in Domestic Sales (Barcelona)



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

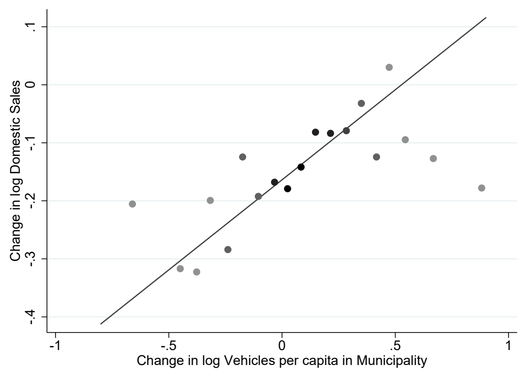
Notes: Panel (a) illustrates the standardized percentage change in average firm-level domestic sales between the period 2002-2008 and the period 2009-2013. Therefore, if this variable takes any given value p for a given zip code, it means that the average firm located in this zip code experienced a relative 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".

C.5 First-Stage and Reduced-Form Relationships

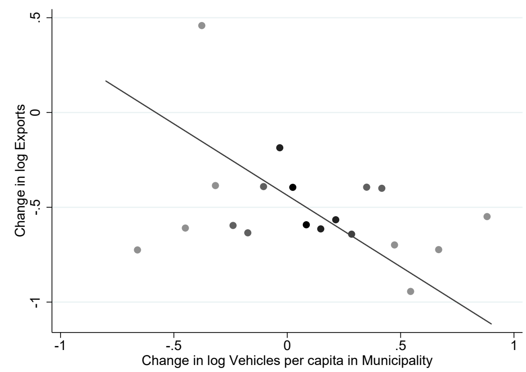
The two panels in Figure C.5 provide a graphical representation of the relationship between our baseline instruments (i.e., the local instrument, in panel A, and the gravity-based instrument, in panel B) and the boom-to-bust change in both the log of domestic sales (left figures) and exports (right figures). 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.

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

(a) Panel A: Local Instrument

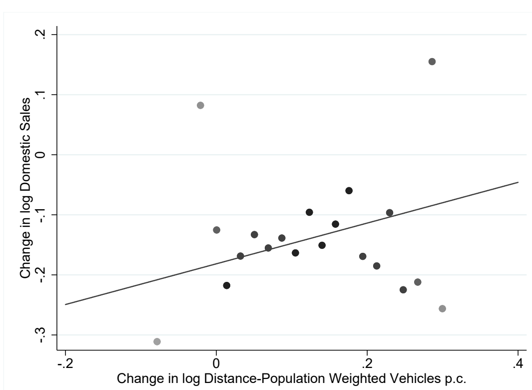


First-Stage Relationship

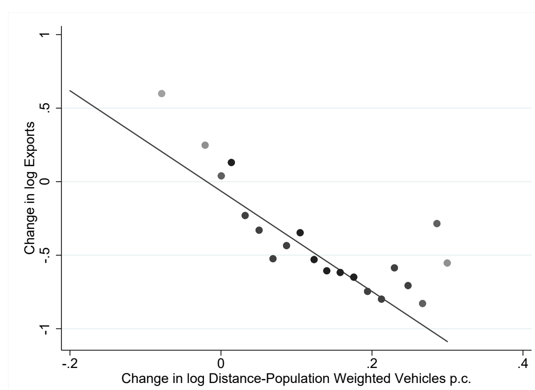


Reduced-Form Relationship

(b) Panel B: Gravity-based Instrument



First-Stage Relationship



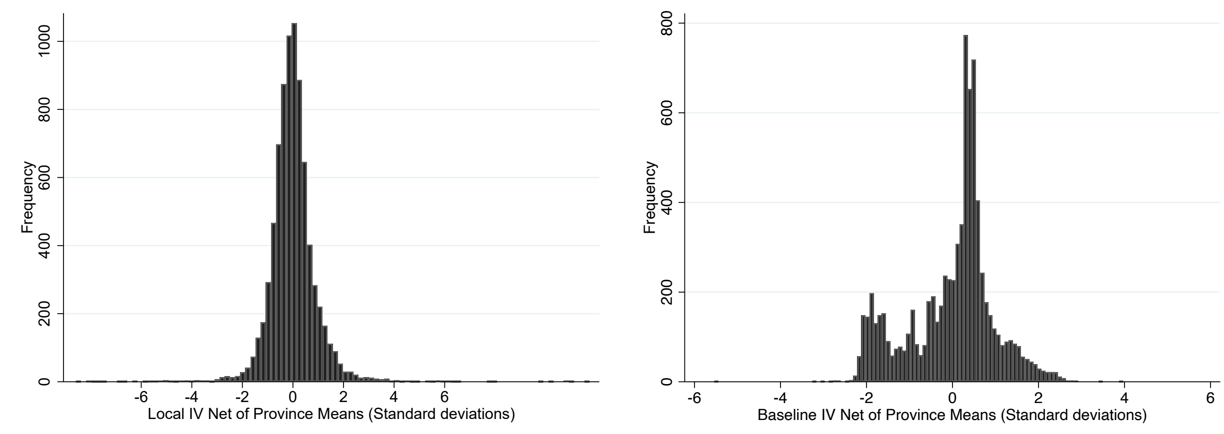
Reduced-Form Relationship

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.

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.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.

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.

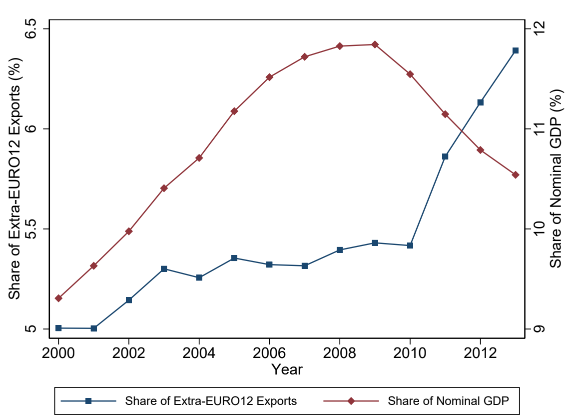
Each of the figures in this Appendix contains two panels. Panel (a) plots, relative to the total for all EMU-12 countries, a country's share of goods exports to non-EMU-12 countries and its share of nominal GDP. Panel (b) includes an analogous plot but excludes from the export data all export flows in 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, and the nominal GDP data is from the AMECO database. 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

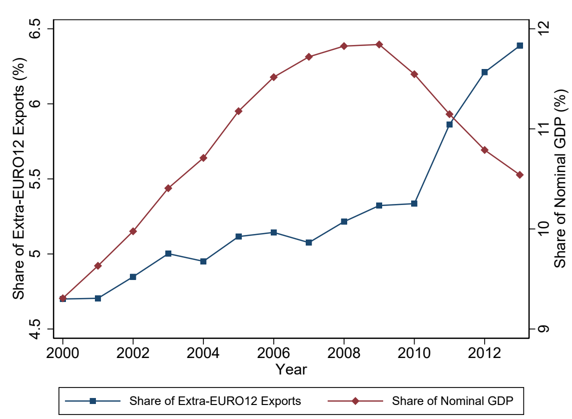
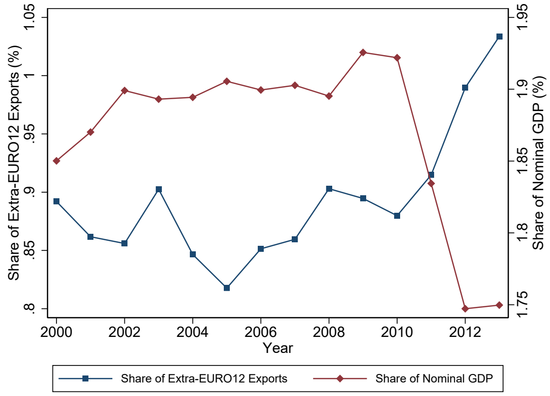


Figure D.2: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Portugal

(a) All Industries



(b) Excluding Motor Vehicle Industry

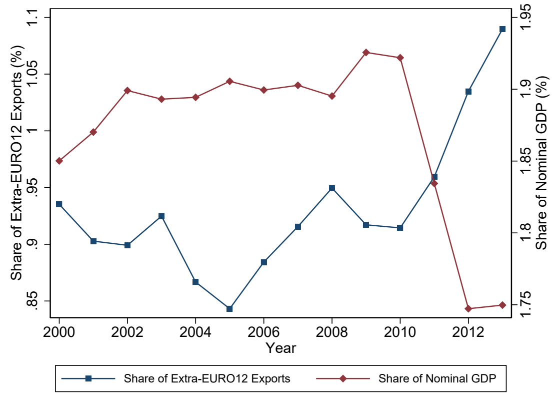
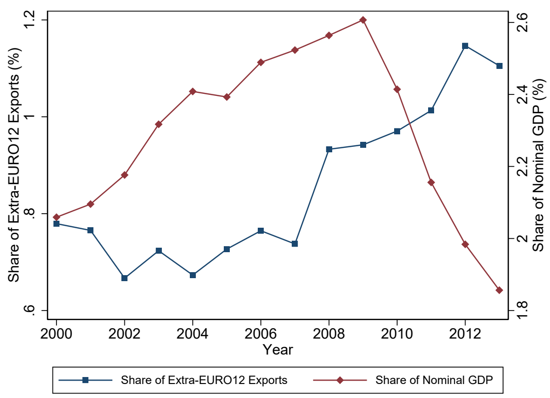


Figure D.3: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Greece

(a) All Industries



(b) Excluding Motor Vehicle Industry

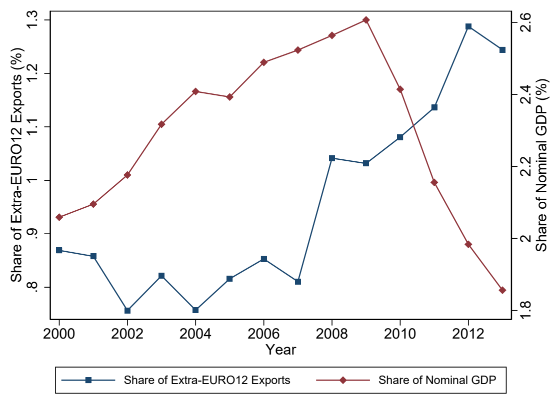
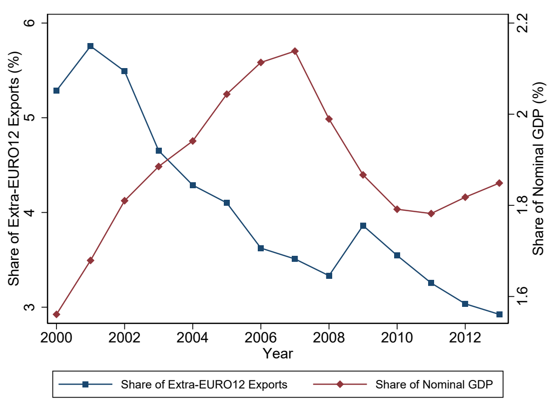


Figure D.4: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Ireland

(a) All Industries



(b) Excluding Motor Vehicle Industry

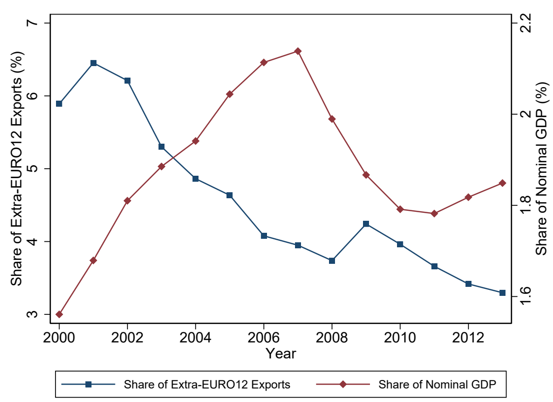


Figure D.5: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Italy

(a) All Industries

(b) Excluding Motor Vehicle Industry

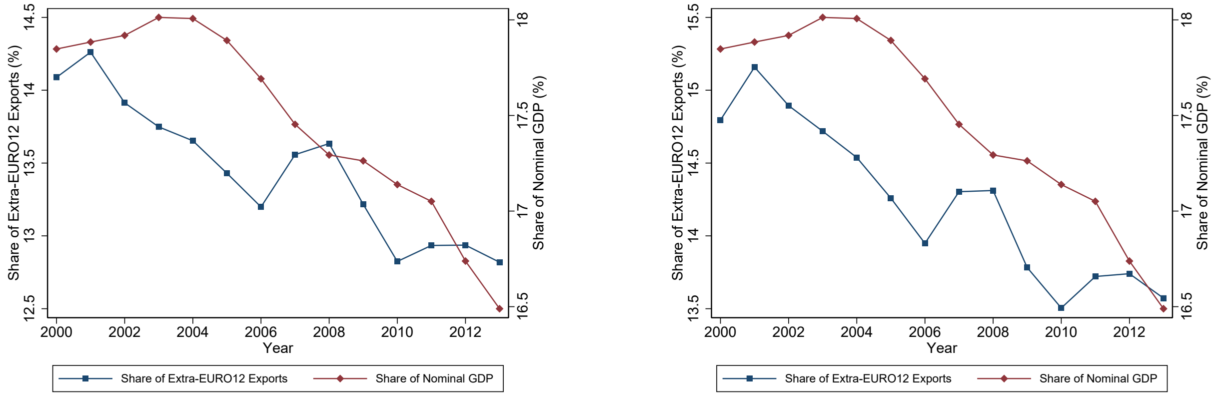


Figure D.6: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Germany

(a) All Industries

(b) Excluding Motor Vehicle Industry

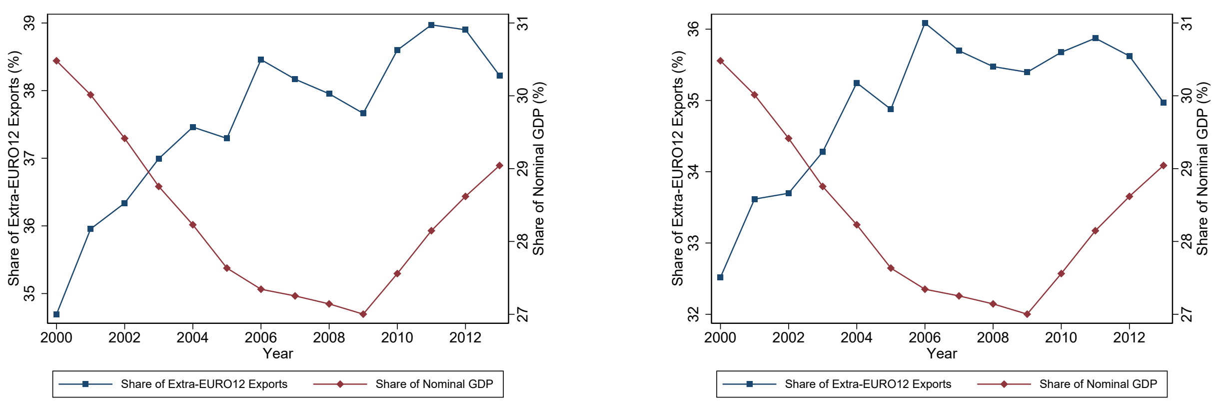


Figure D.7: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: France

(a) All Industries

(b) Excluding Motor Vehicle Industry

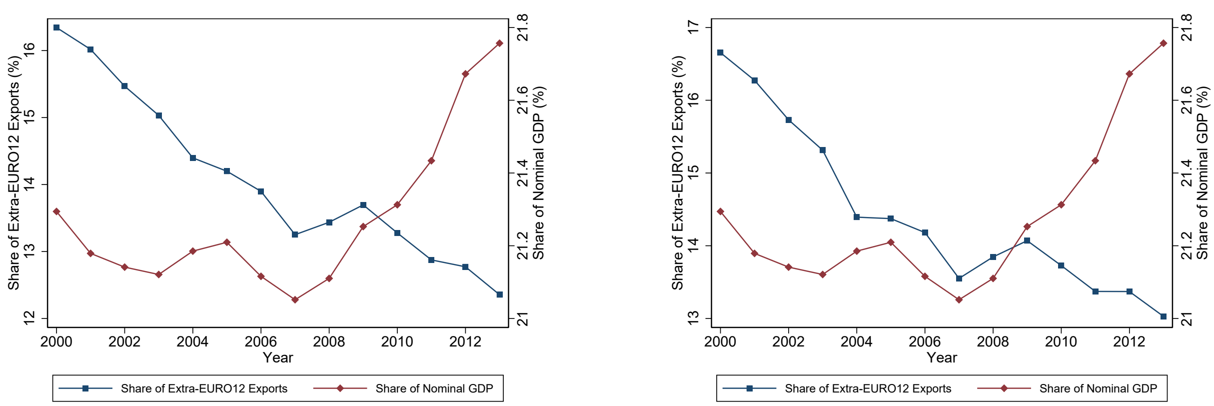
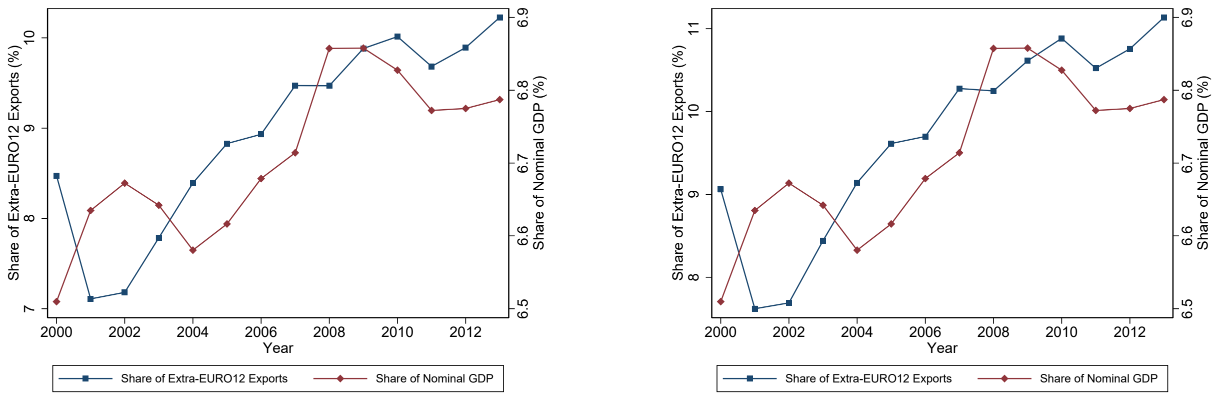


Figure D.8: Share of Extra-Eurozone Exports and GDP in the EMU-12 Countries: Netherlands

(a) All Industries

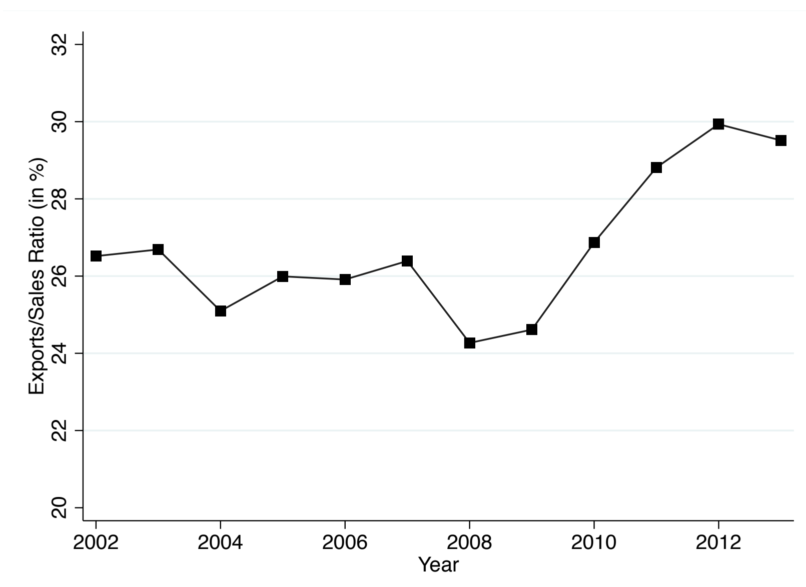
(b) Excluding Motor Vehicle Industry



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.

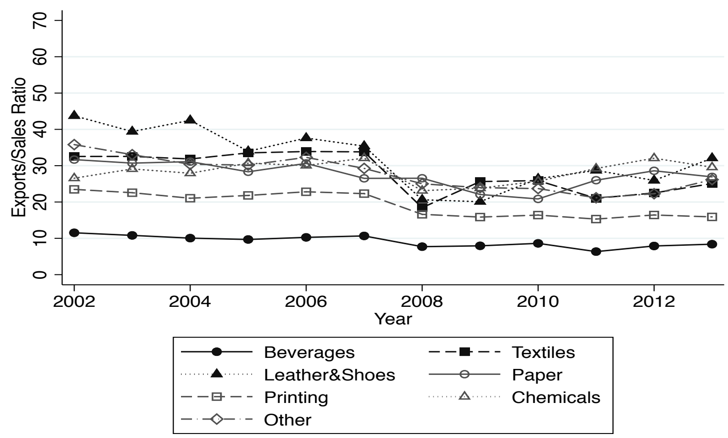
Figure D.9: Exports-to-Sales Ratio for Continuing Exporters (Baseline Sample)



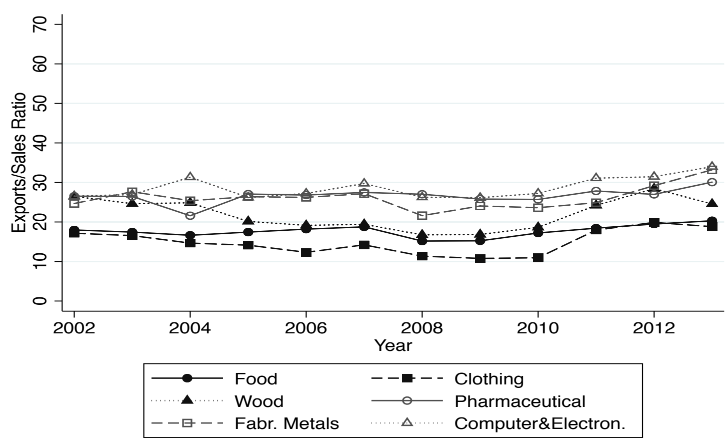
There is considerable variation in the exports-to-sales ratio across sectors. We divide the 21 subsectors of manufacturing considered in our analysis (note that we always exclude tobacco, oil refining and motor vehicles) into three groups depending on whether they had negative, low (positive but below 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

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

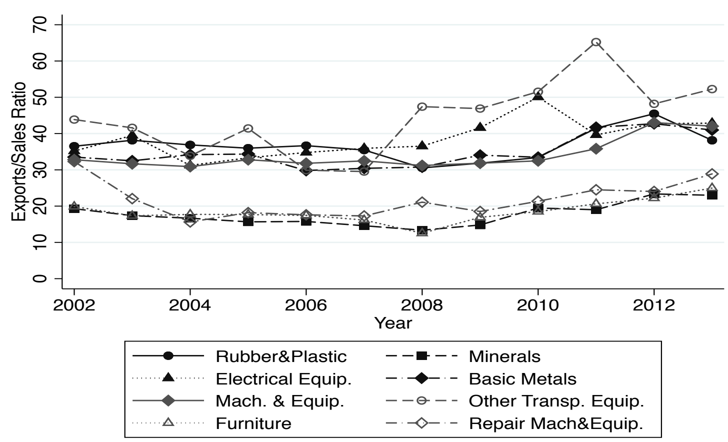
(a) Negative Growth Sectors



(b) Low Growth Sectors



(c) High Growth Sectors



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.

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 U.N. Comtrade 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:

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

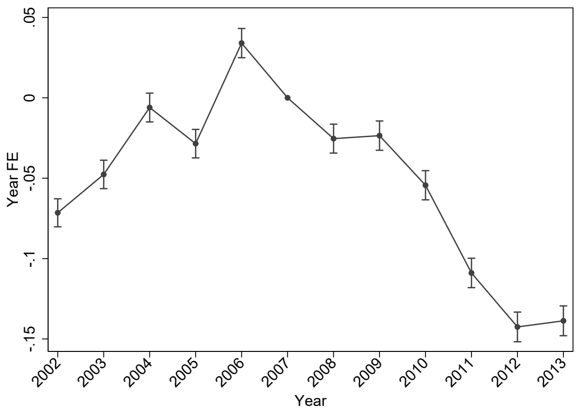
where ω_{jdst} equals the ratio of exports of good s from country j to destination d in year t relative to the total exports of good s from all eleven countries we have selected as comparison group.

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.

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

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

(a) With Constant Weights



(b) With Time-Varying Weights

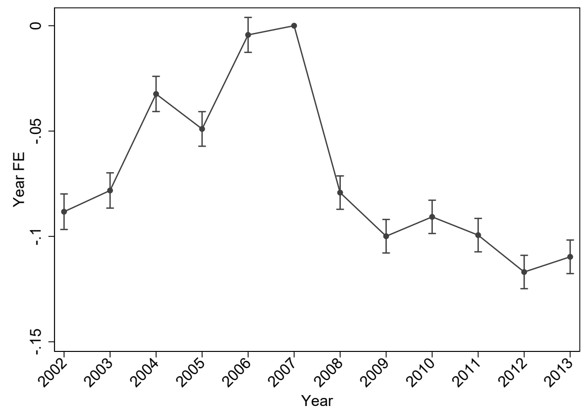
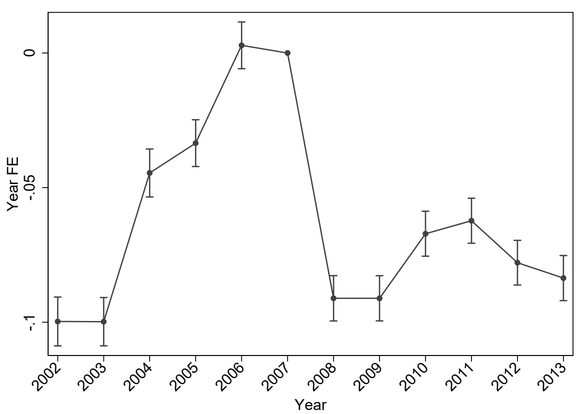
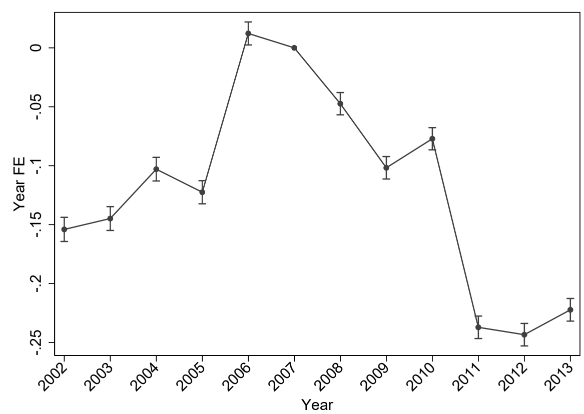


Figure D.12: Export Unit Values of Spain Relative to Germany and France

(a) Relative to Germany



(b) Relative to France



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.

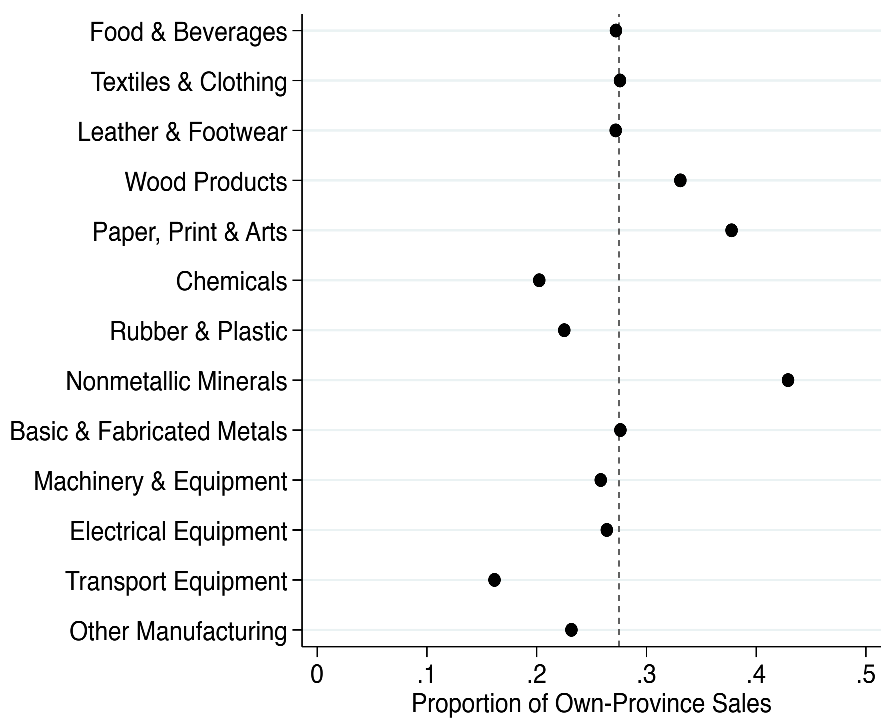
D.4 Home Bias in Firms' Tax Records versus the C-Intereg Dataset

In this Appendix, we briefly compare some aggregate statistics on firms' within-Spain sales from our tax records and from the C-Intereg dataset. 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, for 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 et al., 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.,

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

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.,

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

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

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

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 7 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

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

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

$$\pi_{ijt} = \frac{1}{\sigma} R_{ijt}. \tag{E.11}$$

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 7.1 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 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

[3] In our empirical application (see section 7.1), 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.

$t = 1$ 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 $\Delta\nu_{it}^{x}$ 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_ξ^2 (in column 2); the value of σ_f^2 (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 all firms export (i.e., whenever we set to zero the fixed costs of selling in one of the foreign markets), the downward bias affecting the two IV estimators considered in Table E.1 is very small; typically, the average and median point estimates are within one standard deviation of the true parameter value. 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, and this bias increases in the variance of the demand shock, σ_ξ^2, and decreases in the variance of the fixed costs of selling to a market, σ_f^2. Fifth, for all parameterizations we consider, the distribution of the OLS estimator as well as the distributions of the two IV estimators vary very little as we change the number of countries J.

[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.	0.218	-2.373	-2.373
1	Yes	1	1	10	*Med.*	0.218	-2.370	-2.370
					Std. dev.	0.045	0.075	0.075
					Avg.	0.489	-3.785	-2.111
2	No	1	1	10	*Med.*	0.489	-3.447	-2.029
					Std. dev.	0.213	1.863	0.953
					Avg.	0.630	-2.411	-2.524
3	Yes	1	5	10	*Med.*	0.632	-2.405	-2.519
					Std. dev.	0.046	0.160	0.141
					Avg.	0.774	-2.434	-2.512
4	No	1	5	10	*Med.*	0.776	-2.433	-2.518
					Std. dev.	0.056	0.191	0.182
					Avg.	0.287	-2.406	-2.995
5	Yes	5	1	10	*Med.*	0.287	-2.403	-2.993
					Std. dev.	0.041	0.094	0.121
					Avg.	0.055	-4.112	-3.887
6	No	5	1	10	*Med.*	0.053	-4.106	-3.889
					Std. dev.	0.094	0.257	0.309
					Avg.	0.216	-2.543	-2.806
7	Yes	1	1	14	*Med.*	0.214	-2.546	-2.806
					Std. dev.	0.046	0.194	0.116
					Avg.	0.465	-3.756	-2.304
8	No	1	1	14	*Med.*	0.456	-3.531	-2.226
					Std. dev.	0.186	1.422	0.836
					Avg.	0.693	-2.388	-2.502
9	Yes	1	5	14	*Med.*	0.695	-2.387	-2.494
					Std. dev.	0.048	0.165	0.151
					Avg.	0.880	-2.424	-2.538
10	No	1	5	14	*Med.*	0.883	-2.416	-2.529
					Std. dev.	0.054	0.190	0.176
					Avg.	0.294	-2.603	-3.311
11	Yes	5	1	14	*Med.*	0.294	-2.599	-3.315
					Std. dev.	0.044	0.104	0.147
					Avg.	0.082	-4.060	-4.093
12	No	5	1	14	*Med.*	0.077	-4.064	-4.076
					Std. dev.	0.084	0.217	0.285

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).

E.3 Biases in the Extensive Margin of Exports

We extend here the analysis in section 2 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

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

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 2 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 2, 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

$$plim(\hat{\beta}_{IV}) = \frac{cov(u^{\xi}_{ix} + u^{\varphi}_{i} - u^{\omega}_{i} - \frac{\sigma}{\sigma-1}u^{F}_{i}, \mathcal{Z}_{id})}{cov(u^{\xi}_{id} + u^{\varphi}_{i} - u^{\omega}_{i}, \mathcal{Z}_{id})}. \quad \text{(E.22)}$$

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 2, 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 2.1 (see discussion of equation (12)) and in section 7.1, 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 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	HS-4	HS-6
	(1)	(2)	(3)
Δ Ln(Vehicles p.c.)	-2.841	-0.622	-0.188
	(2.189)	(0.469)	(0.177)
	(2.340)	(0.504)	(0.191)
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 et al., 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 2) or increasing (as assumed in section 7).

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 Antràs et al. (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

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

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

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

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

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

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

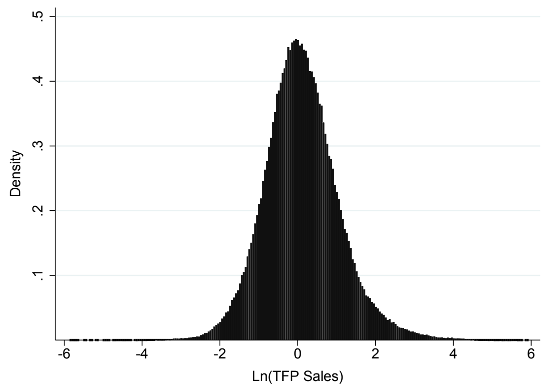
We summarize here the production function and productivity estimates that we obtain both when we assume a production function that is Leontief in materials (see Appendix F.1 for the corresponding estimation approach) and when we assume instead a production function that is Cobb-Douglas in materials (see 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.

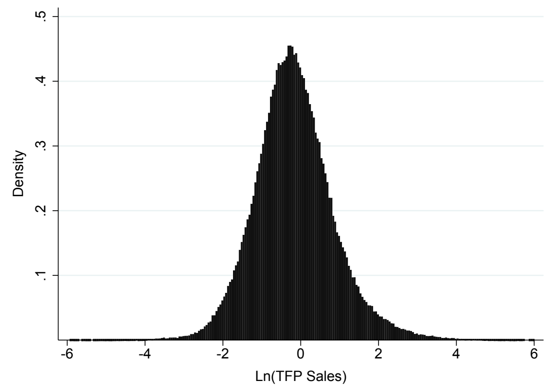
Under the assumption that the production function is Cobb-Douglas in materials, we obtain the following estimates. In the boom period, the average elasticities of value added with respect to labor and capital are 0.76 and 0.19, respectively; the average annual autocorrelation in performance is 0.78;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

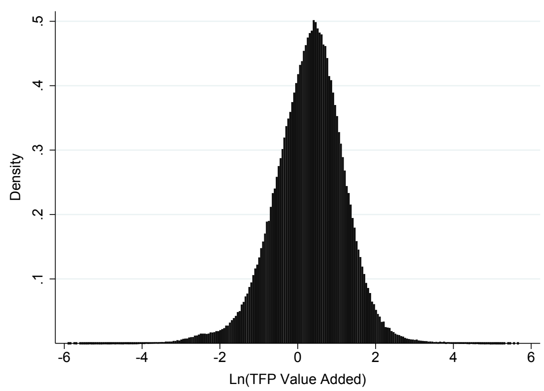
Figure F.1: Productivity Estimates



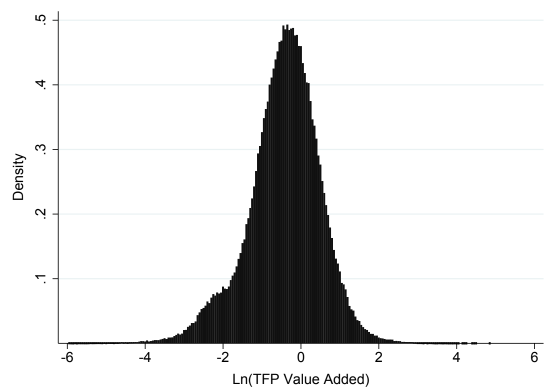
(a) Density of Ln(TFP Sales) in Boom



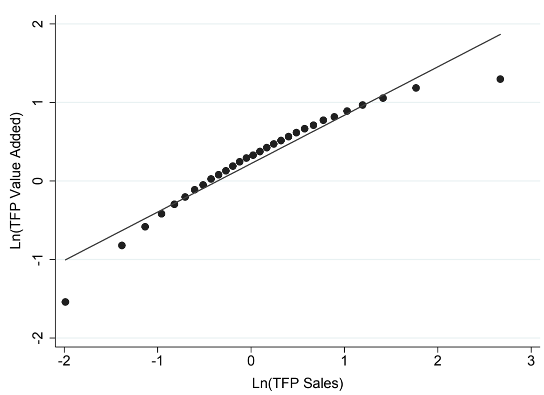
(b) Density of Ln(TFP Sales) in Bust



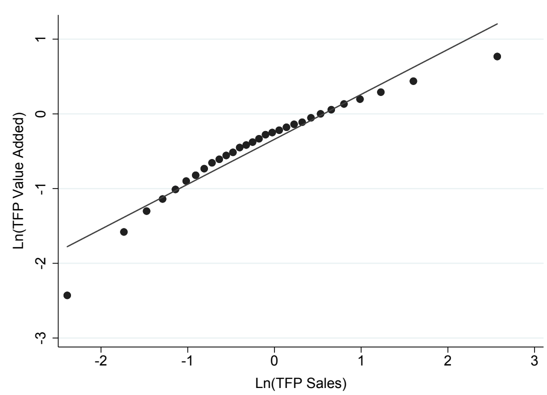
(c) Density of Ln(TFP Value Added) in Boom



(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.

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 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 8. 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 close to log-normally distributed. 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, on average, 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}	-0.143	0.008^{a}	-0.011
		(0.005)	(0.108)	(0.003)	(0.057)
Ln(Vehicles p.c. in municipality)	0.191^{a}				
	(0.044)				
Ln(TFP)	1.140^{a}	0.068^{a}	0.254^{b}	0.062^{a}	0.082
	(0.014)	(0.007)	(0.123)	(0.004)	(0.065)
Ln(Average Wages)	-0.568^{a}	-0.046^{a}	-0.139^{b}	-0.041^{a}	-0.051
	(0.013)	(0.006)	(0.061)	(0.004)	(0.032)
Observations	125,054	125,054	125,054	125,054	125,054
R-squared	0.983	0.843	-0.040	0.920	0.016
Firm FE	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes
Sector-Period FE	Yes	Yes	Yes	Yes	Yes
F-statistic on IV	18.62				
Mean of Dep. Var.		0.171	0.171	0.115	0.115
Ext-Margin Elasticity		0.121	-0.838	0.066	-0.093

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.				-0.232^a			
(munic. level)				(0.053)			
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	29.10	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 6.4 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
ΔLn(Domestic Sales)	-0.284^a	-1.327^a	0.027	-0.686^c
	(0.030)	(0.337)	(0.028)	(0.387)
ΔLn(Average Wages)	-0.712^a	-1.240^a	-0.749^a	-0.925^a
	(0.059)	(0.178)	(0.068)	(0.115)
ΔLn(TFP Sales): Baseline	1.522^a	2.533^a		
	(0.051)	(0.323)		
ΔLn(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.

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

	Panel A: Local Instrument							
Instrument:	Including lags of IV				Combined lags of IV			
	1st Stage (1)	2SLS (2)	1st Stage (3)	2SLS (4)	1st Stage (5)	2SLS (6)	1st Stage (7)	2SLS (8)
Ln(Domestic Sales)		0.256		0.305		1.826		3.142
		(0.823)		(0.777)		(2.647)		(4.749)
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 $t-1$					0.041			
					(0.048)			
Average of Ln(Vehicles p.c. in municipality) in $t-1$ and $t-2$							0.035	
							(0.047)	
Ln(TFP)	0.865^a	0.933	0.863^a	0.909	0.920^a	-0.548	0.920^a	-1.757
	(0.040)	(0.716)	(0.050)	(0.673)	(0.030)	(2.435)	(0.030)	(4.375)
Ln(Average Wages)	-0.344^a	-0.437	-0.350^a	-0.448	-0.372^a	0.200	-0.372^a	0.690
	(0.035)	(0.290)	(0.044)	(0.278)	(0.029)	(0.986)	(0.029)	(1.776)
Observations	45,384	45,384	35,863	35,863	59,951	59,951	59,951	59,951
F-statistic	2.55		1.95		0.74		0.57	

	Panel B: Gravity-based Instrument							
Instrument:	Including lags of IV				Combined lags of IV			
	1st Stage (1)	2SLS (2)	1st Stage (3)	2SLS (4)	1st Stage (5)	2SLS (6)	1st Stage (7)	2SLS (8)
Ln(Domestic Sales)		1.377		1.606		1.757		2.028
		(1.300)		(2.089)		(1.447)		(1.642)
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 $t-1$					0.705			
					(0.466)			
Average of Ln(Dist-Pop-Weighted Vehicles p.c.) in $t-1$ and $t-2$							0.686	
							(0.451)	
Ln(TFP)	0.864^a	-0.035	0.862^a	-0.213	0.920^a	-0.488	0.920^a	-0.738
	(0.042)	(1.136)	(0.052)	(1.817)	(0.027)	(1.358)	(0.027)	(1.539)
Ln(Average Wages)	-0.344^a	-0.052	-0.350^a	0.009	-0.373^a	0.178	-0.373^a	0.279
	(0.043)	(0.453)	(0.050)	(0.746)	(0.032)	(0.530)	(0.032)	(0.599)
Observations	45,384	45,384	35,863	35,863	59,954	59,954	59,954	59,954
F-statistic	1.16		0.71		2.29		2.32	

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.

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 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 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. Multi-nationals (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 Vehicles p.c.)	1.312^a		1.394^a		1.457^a	
	(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 Vehicles p.c.)	0.888^a		0.913^a		0.962^a	
	(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.

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.

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 (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

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.

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).

Table G.12 reports the results of estimating our main intensive-margin regression using the 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 (1)	Ln(Avg. Empl.) 2002-2008 (2)	Ln(Avg. Assets) 2002-2008 (3)	Exports-to-Sales 2002-2008 (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

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.

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.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 1.44 and 1.24, 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.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	-1.037^a	-0.920^a
	(0.050)	(0.049)	(0.049)	(0.049)
Ln(Population)	1.346^a	1.332^a	1.339^a	1.327^a
	(0.028)	(0.028)	(0.024)	(0.023)
Dummy for own-province flows		1.449^a		1.239^a
		(0.150)		(0.130)
Observations	2,597	2,597	2,541	2,541
R-squared	0.83	0.84	0.84	0.84
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)					-10.281^a	-2.093^a	-1.759^a	-1.614^a
					(3.541)	(0.317)	(0.235)	(0.245)
Δ Ln(Dist-Pop-Weighted Vehicles p.c.)	0.338^a	1.210^a	1.364^a	1.334^a				
	(0.121)	(0.145)	(0.135)	(0.119)				
Δ Ln(TFP)		0.829^a	1.032^a	1.024^a		2.633^a	2.884^a	2.817^a
		(0.028)	(0.029)	(0.028)		(0.239)	(0.220)	(0.211)
Δ Ln(Average Wages)			-0.622^a	-0.526^a			-1.625^a	-1.390^a
			(0.037)	(0.047)			(0.174)	(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	7.82	69.24	102.66	126.55				

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.

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.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)					-3.077[a]	-1.228[b]	-1.226[a]	-0.986[c]
					(1.162)	(0.527)	(0.466)	(0.587)
Δ Ln(Dist-Pop-Weighted Vehicles p.c.)	0.134[a]	0.235[a]	0.235[a]	0.207[a]				
	(0.046)	(0.059)	(0.061)	(0.058)				
Δ Ln(TFP)		0.796[a]	0.987[a]	0.974[a]		1.949[a]	2.361[a]	2.208[a]
		(0.026)	(0.028)	(0.028)		(0.401)	(0.426)	(0.527)
Δ Ln(Average Wages)			-0.603[a]	-0.505[a]			-1.304[a]	-1.073[a]
			(0.038)	(0.050)			(0.263)	(0.269)
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	8.51	15.93	14.66	12.86				

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 $\beta_{ownprov} = 1$.

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.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)					-2.597[b]	-1.092[b]	-1.116[b]	-0.866
					(1.127)	(0.583)	(0.529)	(0.676)
Δ Ln(Dist-Pop-Weighted Vehicles p.c.)	0.112[a]	0.184[a]	0.182[a]	0.158[a]				
	(0.037)	(0.046)	(0.047)	(0.044)				
Δ Ln(TFP)		0.795[a]	0.986[a]	0.973[a]		1.841[a]	2.252[a]	2.092[a]
		(0.026)	(0.028)	(0.028)		(0.446)	(0.490)	(0.615)
Δ Ln(Average Wages)			-0.603[a]	-0.505[a]			-1.237[a]	-1.012[a]
			(0.039)	(0.050)			(0.300)	(0.315)
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	8.90	16.23	14.80	12.60				

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)	0.477^a	0.479^a	0.491^a	-0.098^b
	(0.048)	(0.047)	(0.047)	(0.043)
ΔLn(TFP)				1.371^a
				(0.048)
ΔLn(Average Wages)				-0.512^a
				(0.095)
Observations	4,488	4,488	4,488	4,488
R-squared	0.108	0.128	0.162	0.308
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)					18.554	19.767	18.324	-1.025^a
					(24.501)	(26.415)	(22.459)	(0.296)
ΔLn(Dist-Pop-Weighted Vehicles p.c.)	-0.188	-0.201	-0.209	1.796^a				
	(0.258)	(0.275)	(0.263)	(0.182)				
ΔLn(TFP)				1.230^a				2.482^a
				(0.026)				(0.323)
ΔLn(Avg. Wages)				-0.401^a				-0.868^a
				(0.056)				(0.144)
Observations	4,488	4,488	4,488	4,488	4,488	4,488	4,488	4,488
Province FE	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Sector FE	No	No	Yes	Yes	No	No	Yes	Yes
F-statistic	0.53	0.53	0.63	97.46				

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 et al. (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 et al. (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.

Contrary to the findings in Berman et al. (2015), and consistently with the estimates we present in section 5 in the main draft, our OLS estimates of the elasticity domestic sales with respect to

[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 et al. (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.

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 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 7.1 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 et al. (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 et al. (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 et al. (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.

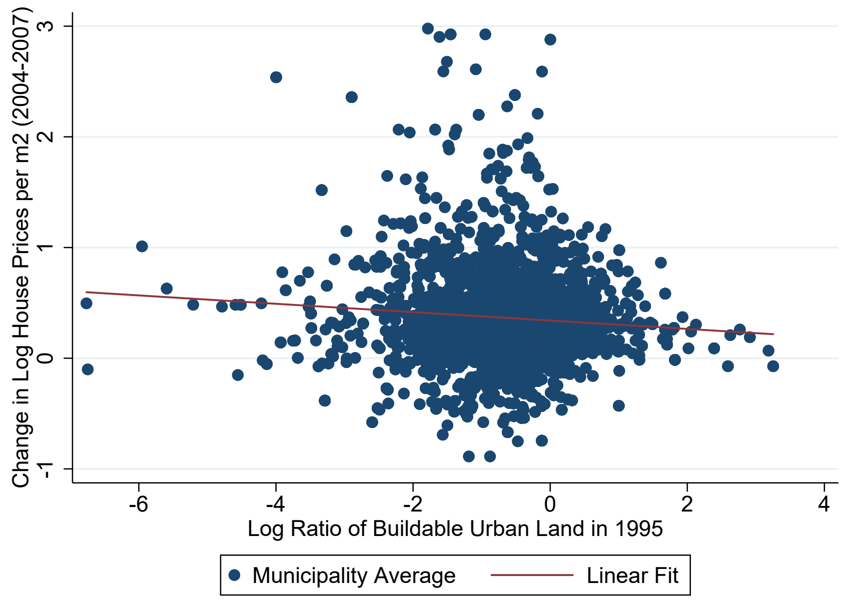
As described in section 3.1 of the main text, the Great Recession in Spain was largely driven by a real estate bubble. Our first alternative instrument thus attempts to identify an exogenous source of the intensity of the bubble across different locations. More precisely, we construct ratios of available 'buildable' urban land to urban land with already built structures in the year 1995 (a year sufficiently removed from the housing boom). 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 2002 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 2002 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

[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



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 2002 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 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

[8]The relevance and validity of our instrument does not depend on the fact that we multiply the municipality-specific 2002 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 2002 construction wage bill. See Goldsmith-Pinkham et al. (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)		
(Weighted by Distance and Population)					
ΔLn(construction employment) ×				0.428^a	
2001 empl. share in munic.				(0.074)	
(Weighted by Distance and Population)					
ΔLn(construction turnover) ×					0.160^a
2001 turnover share in munic.					(0.025)
(Weighted by Distance and Population)					
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.

particular, we use the measure in Basco et al. (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). We calculate the market value price per square meter for each residential housing transaction and 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 (Instituto Nacional de Estadística). 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	1st Stage	2SLS	OLS	1st Stage	2SLS	OLS	1st Stage	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(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	1st Stage	2SLS	OLS	1st Stage	2SLS	OLS	1st Stage	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(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

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)		0.068^a	-0.101^a	0.050^a	0.041^b
		(0.007)	(0.035)	(0.006)	(0.019)
Ln(Dist-Pop-Weighted Vehicles p.c.)	1.007^a				
	(0.108)				
Ln(TFP)	1.184^a	0.013^b	0.208^a	0.012^a	0.023
	(0.016)	(0.006)	(0.041)	(0.004)	(0.020)
Ln(Average Wages)	-0.599^a	-0.019^a	-0.116^a	-0.016^a	-0.022^b
	(0.016)	(0.006)	(0.022)	(0.003)	(0.010)
Observations	125,054	125,054	125,054	125,054	125,054
F-statistic	87.13				
Mean of Dep. Var.		0.171	0.171	0.115	0.115
Ext-Margin Elasticity		0.395	-0.591	0.439	0.358

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.

and own-province dummies in the gravity equation we use to compute the estimated distance

Table I.5: Alternative Instruments

Dependent Variable:	ΔLn(Total Sales)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Gravity: mun-mun flows (Baseline)	0.888^a (0.103)						
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Baseline incl. own mun. & prov. dummies		0.901^a (0.103)					
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Gravity: distance dummies			0.632^a (0.070)				
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Gravity: firm-mun flows				0.881^a (0.089)			
ΔLn(Weighted Vehicles p.c.) Weights: firm-level mun. shares					0.257^a (0.069)		
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Fixed coefficients: $\beta_{pop} = 1$, $\beta_{dist} = -1$						0.243^a (0.070)	
ΔLn(Dist-Pop-Wght. Vehicles p.c.) Baseline in levels Adão et al. (2019) std. error							0.651^a (0.083) (0.099)
F-statistic	75.00	75.81	80.65	98.90	13.86	12.16	61.19 43.13
	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Total Sales)	-2.374^a (0.526)	-2.390^a (0.523)	-2.018^a (0.463)	-2.428^a (0.447)	-4.374^a (1.266)	-2.081^b (0.854)	-2.526^a (0.579) (0.610)

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.

elasticity that enters in the construction of the instrument weights. In column 3, we use a more 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 (firm level)		-0.404^a					
		(0.135)					
ΔShare of Temp. Workers (munic. level)			-0.094				
			(0.188)				
ΔManufacturing Empl. p.c. (munic. level)				-0.384^a			
				(0.068)			
ΔLn(Financial Costs) (firm level)					-0.061^a		
					(0.019)		
Financial Costs in Boom (firm level)						-0.014	
						(0.019)	
Financial Costs in Boom (munic. level)							-0.063
							(0.047)
F-Statistic	75.00	74.48	73.55	74.33	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

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, ΔLn(X) is the difference in Ln(X) between its average in the 2009-2013 period and its average in the 2002-2008 period. All specifications include sector fixed effects.

short-run constant marginal costs according to this micro-foundation.

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) $\times$ High	1.434			0.462		
Output elasticity wrt Materials	(1.358)			(0.782)		
ΔLn(Total Sales) $\times$ High		1.335			2.205^{c}	
Output elasticity wrt Labor		(1.263)			(1.143)	
ΔLn(Total Sales) $\times$ (High			3.037			3.963^{a}
Output elast. wrt Labor			(2.364)			(1.350)
$\times$ Temp. Ratio						
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, ΔLn(X) is the difference in 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 8.

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

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

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:

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

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

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

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_i^{φ} and u_i^{ω} and the export demand shocks u_{ix}^{ξ}.

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

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

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.,

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

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 8). 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 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 $(\sigma - 1)\lambda/(1 + \lambda)$.

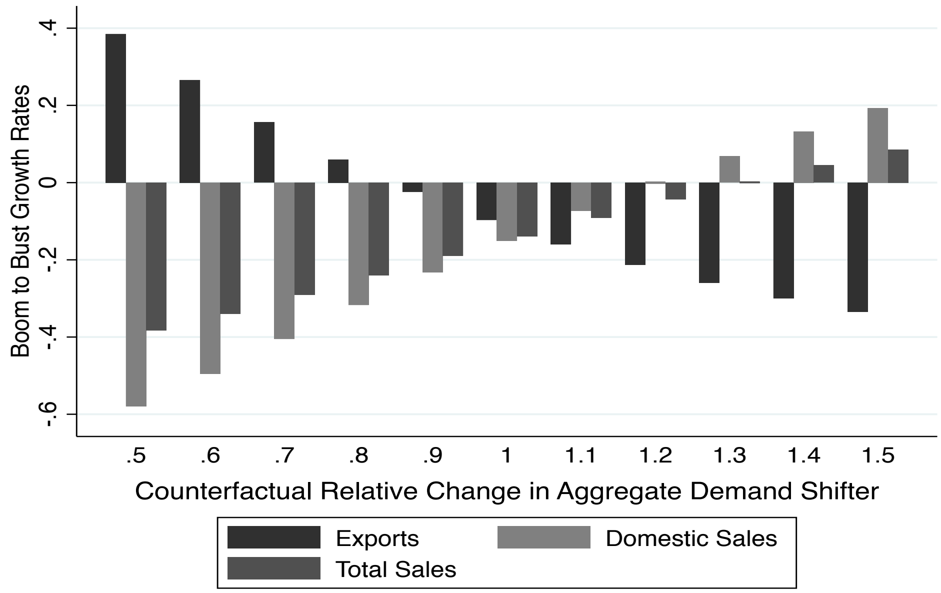
To quantify the impact that the value of the parameter $(\sigma - 1)\lambda/(1 + \lambda)$ has on our results, we recompute the baseline counterfactual exercise for several different values of this parameter.

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.

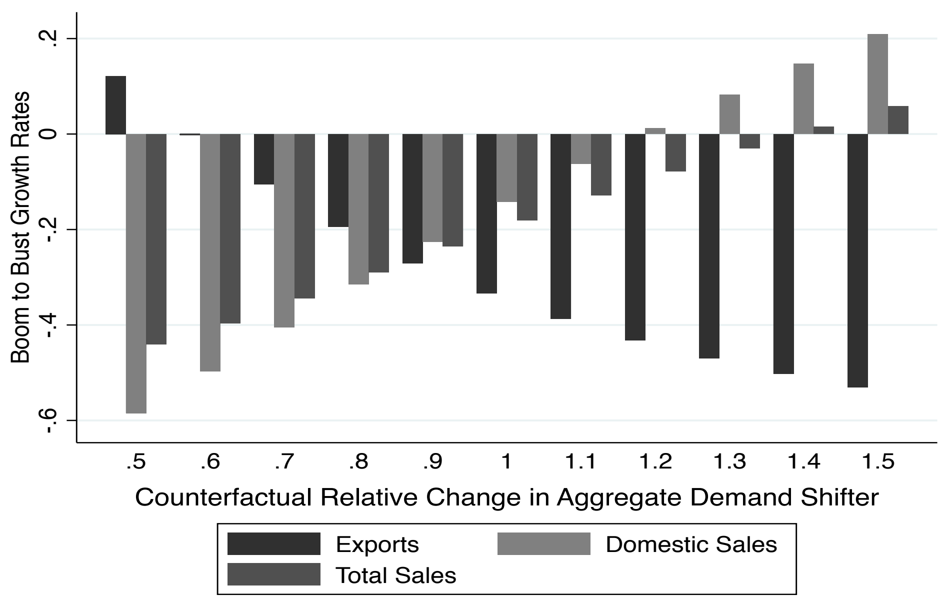
In Table J.1, we present results on the implied contribution of the vent-for-surplus mechanism to the boom-to-bust growth in exports for several additional values of the parameter $(\sigma-1)\lambda/(1+\lambda)$.

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

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



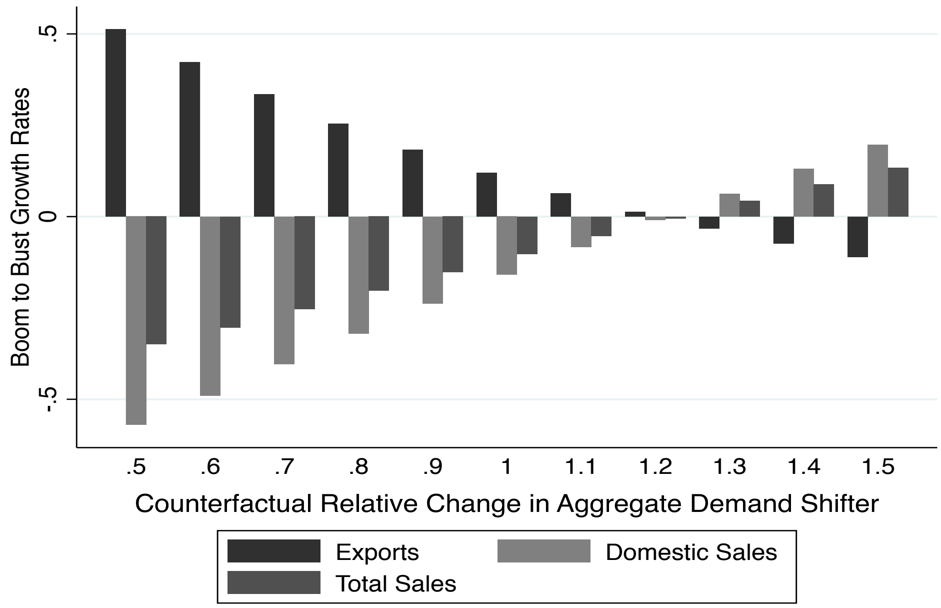
(b) With a 25% Increase in Trade Costs ($\Gamma_\tau = 1.25$)



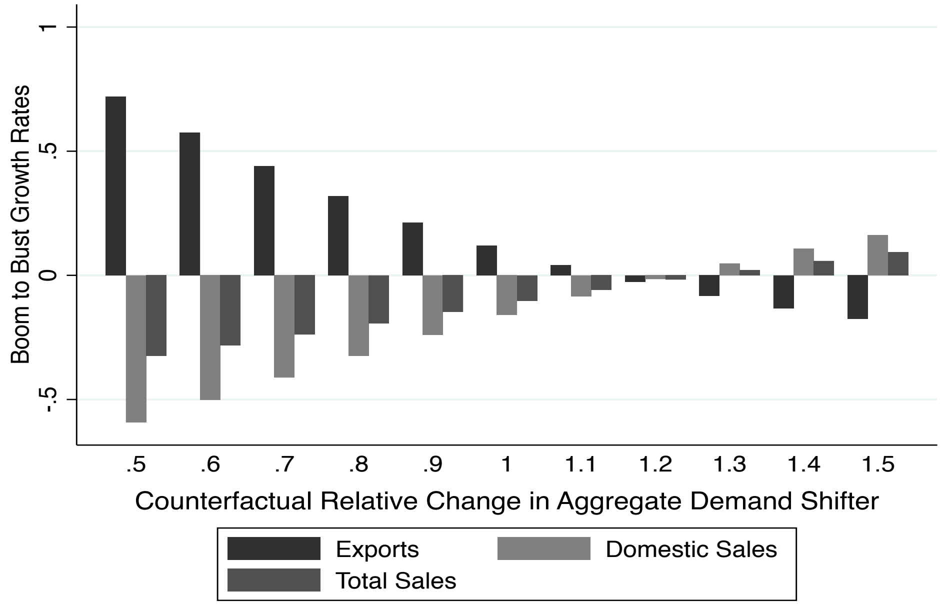
Notes: The horizontal axis indicates the value of Γ_B. The export and domestic sales growth rates indicated in the vertical axis correspond to those predicted by equations (J.5), (J.11), (J.18) and (J.19). Panel (a) imposes $\Gamma_\tau = 1.1$; Panel (b) imposes $\Gamma_\tau = 1.25$. Given these counterfactual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

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
$0.50 \times 2.374 = 1.1870$	31.94%
$0.75 \times 2.374 = 1.7805$	42.79%
2.374	51.71%
$1.25 \times 2.374 = 2.9675$	59.38%
$1.50 \times 2.374 = 3.5610$	66.06%
$2 \times 2.374 = 4.7480$	76.98%

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%).

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

$$\begin{aligned}\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] &= \ln\left[\frac{B_{sx1}}{B_{sx0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right] + \frac{(\sigma-1)}{1+\lambda}\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - (\sigma-1)\ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right] \\ &\quad + \sigma\ln\left[\frac{P_{sx1}}{P_{sx0}}\right] - \frac{(\sigma-1)}{1+\lambda}\left(\lambda + \frac{\psi}{1+\psi}\right)\ln\left[\frac{R'_{i1}}{R_{i0}}\right], \end{aligned} \tag{J.28a}$$

$$\begin{aligned}\ln\left[\frac{R'_{id1}}{R_{id0}}\right] &= \ln\left[\frac{B'_{sd1}}{B_{sd0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right] + \frac{(\sigma-1)}{1+\lambda}\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - (\sigma-1)\ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right] \\ &\quad + \sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\right] - \frac{(\sigma-1)}{1+\lambda}\left(\lambda + \frac{\psi}{1+\psi}\right)\ln\left[\frac{R'_{i1}}{R_{i0}}\right]. \end{aligned} \tag{J.28b}$$

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

$$\begin{aligned}\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] &= \ln\left[\frac{R_{ix1}}{R_{ix0}}\right] \\ &\quad - \frac{(\sigma-1)}{1+\lambda}\left(\lambda + \frac{\psi}{1+\psi}\right)\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \end{aligned} \tag{J.29a}$$

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

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 8) 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.

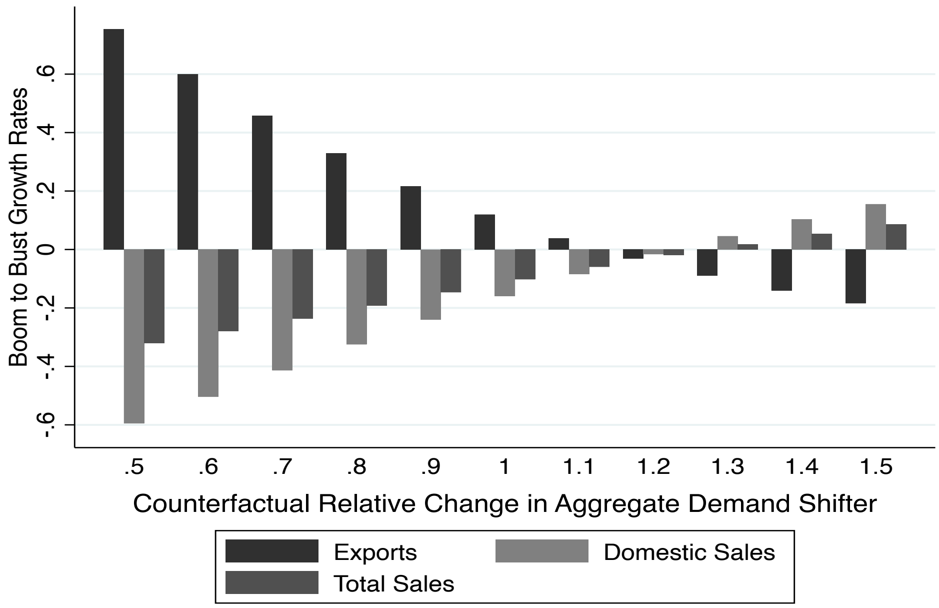
A comparison of the results in Figure 4 (see section 8 in the main text) with those in Figure J.3 reveal that lower elasticities of labor supply imply that firms' exports react more to changes in domestic demand shifters. Thus, the lower the elasticity of the labor supply function that a firm faces, 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 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 8 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 $1 - 1/1.1 = 9.91\%$).

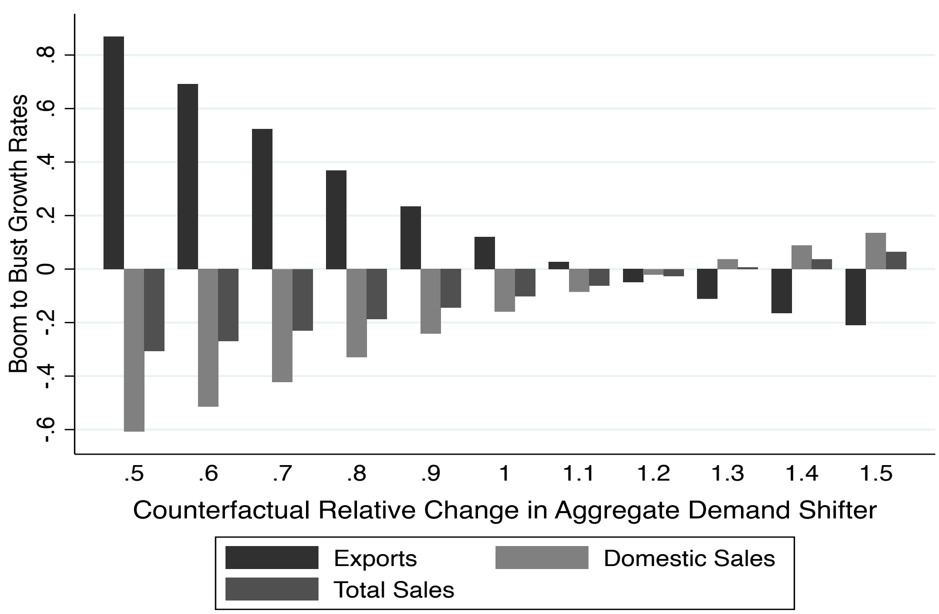
For the corresponding calibrated values of Γ_B, our baseline results with fully elastic labor supply predict a growth in aggregate exports of 5.79%, those that assume a firm-specific labor supply with unit elasticity predict a growth in aggregate exports of 3.81%, and those that assume a perfectly inelastic 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.

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 $(R'_{ix1}/R_{ix0})\chi_{i0} + (R'_{id1}/R_{id0})(1 - \chi_{i0})$.

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