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 well 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.

Keywords: vent-for-surplus, exports, domestic slump, increasing marginal cost.

JEL classification: F12, F14, F62.

Resumen

En este artículo utilizamos variación geográfica plausiblemente exógena en la reducción de la demanda interna causada por la Gran Recesión en España para documentar la existencia de una robusta relación causal negativa a nivel de empresa entre cambios provocados por la demanda en las ventas internas y en los flujos de exportaciones. Las empresas manufactureras españolas cuyas ventas internas se redujeron en mayor medida durante la crisis experimentaron un mayor incremento en sus flujos de exportación, incluso una vez que se controla por sus determinantes de oferta, tales como sus costes laborales. Esta relación negativa entre cambios en las ventas internas y cambios en los flujos de exportación provocados por cambios de la demanda ilustra la capacidad de los mercados de exportación para compensar el impacto negativo de *shocks* de demanda locales. Los resultados presentados en el artículo se racionalizan a través de un modelo estándar de exportaciones con heterogeneidad de empresas que permite la existencia de costes marginales de producción no constantes. Utilizando una versión de este modelo estimada de manera estructural, concluimos que las respuestas a nivel de empresa a la caída de la demanda interna en España podrían explicar alrededor de la mitad del espectacular incremento de las exportaciones de bienes en España (el denominado «milagro exportador español») en el período 2009-2013.

Palabras clave: salida al exterior, exportaciones, crisis interna, coste marginal creciente.

Códigos JEL: F12, F14, F62.

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). 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%, and the unemployment rate shot up from 9.6% to 26.9%. Portugal and Greece experienced comparably marked domestic 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.[1] 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, 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 (see Appendix C.1).[2]

Figure 1: The Spanish Export Miracle
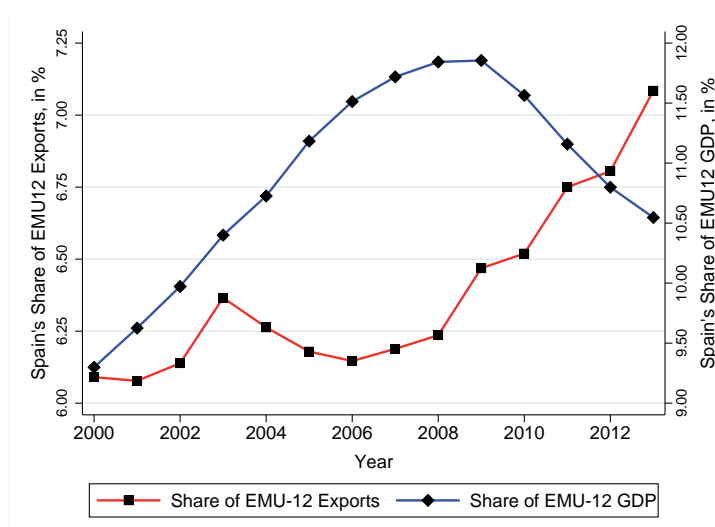
At first glance, this remarkable export performance appears to be consistent with the type of "internal devaluation" processes advocated by international organizations (such as IMF, ECB or the European Commission) since the onset of the crisis. According to this thesis, wage moderation

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[1]The implied 6.9% annual growth in real exports from 2009 to 2013 almost doubled the 3.8% annual growth in real exports during the period 2000-2008.

[2]In Appendix C.1, we replicate Figure 1 for Portugal and Greece, and also for Germany, whose relative GDP increased during the crisis. In all three cases, we observe a negative relationship between these countries' GDP shares in the euro area and their shares in euro area goods exports to other countries. See Section 3 and Appendix B for a description of the data sources underlying these figures.

coupled with a set of structural reforms (most notably labor market reforms) led 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, the adjustment in labor costs achieved via these policies was modest up to 2013 and this channel 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, Portugal and Greece over the period 2010-2013? At least for the case of Spain, an often-invoked alternative explanation relates the growth in exports *directly* to the collapse in domestic demand. According to this hypothesis, the unexpected demand-driven reduction in Spanish firms' domestic sales, in combination with the irreversibility of certain investments in inputs, freed up capacity that these firms used to serve customers abroad.[3] More precisely, this explanation posits that, as domestic demand dropped, 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 firms' exports.[4]

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.[5] 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 product differentiation and economies of scale of the Krugman-Melitz type – assume that firms face constant marginal costs of production, an assumption that implies that firms' domestic and export sales decisions can be studied independently from each other.

In this paper, we leverage Spanish firm-level data from 2002 to 2013, and geographic variation across Spanish regions in the reduction in domestic demand caused by the financial crisis, to study the empirical relevance of the "vent-for-surplus" mechanism. To do so, we divide our sample into a "boom" period (2002-08) and a "bust" period (2009-13), and measure the extent to which, 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. To further isolate demand-driven changes in domestic sales, we exploit the fact that the financial crisis and the Great

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[3]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), a newspaper article which was inspired by an early version of our paper.

[4]Generally, 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 (e.g., reductions in the price of factors or materials, or increases in productivity).

[5]In *The Wealth of Nations* (1776) Book II, Chapter V, Adam 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).

Recession affected different geographical areas in Spain differentially. More specifically, we rely on municipality-level registration data on a major household durable consumption item, vehicles, and use the change in the municipality-level stock of vehicles per capita between 2002-08 and 2009-13 as a proxy for the extent to which the Great Recession affected demand across municipalities. We use this measure of changes in local demand as an instrument for the reduction in the domestic sales of firms located in different parts of Spain.

To understand the properties of our estimates of the causal impact of demand-driven changes in domestic sales on exports, we first base our analysis on a commonly used model of firms' export behavior: a model à la Melitz (2003). For our purposes, this framework serves the role of identifying several empirical challenges that one encounters when measuring the relevance of the "vent-for-surplus" mechanism; i.e., when measuring the causal impact of changes in a firm's domestic sales on exports working exclusively through changes in the firm's domestic demand.[6] We draw three 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 does not reflect a causal impact of the former on the latter. Second, the fact that firm-level domestic sales are computed as the difference between firm-level total sales and exports leads to a non-classical error-in-variables bias that, under plausible conditions, tends to generate a *negative* spurious correlation between exports and domestic sales (see also Berman et al., 2015). Third, an instrumental variable approach that exploits a proxy for 'local demand' as an instrument for the changes in domestic sales of the firms producing in a given locality identifies the causal impact of demand-driven changes in domestic sales on exports as long as it satisfies three conditions: (i) it is indeed a useful proxy for 'local demand' (i.e., the overall propensity to consume of the residents of a locality), (ii) 'local demand' is a good predictor of the domestic sales of Spanish firms producing in a given locality, and (iii) this proxy is not correlated with unobserved covariates that have an independent effect on Spanish firms' exporting decisions (i.e., unobserved marginal cost or export-demand shifters). We discuss each of these three conditions in turn.

Although, given available data, we cannot directly test that "boom-to-bust" changes in the stock of vehicles per capita in the municipality of location of a firm satisfies conditions (i) and (ii), prior work has provided empirical evidence supporting the independent validity of each of these two conditions. First, an extensive literature in empirical macroeconomics has documented that consumption of durable goods (such as vehicles) is strongly procyclical (see, for instance, the survey by Stock and Watson, 1999). Second, a significant impact of highly localized demand shocks on Spain-wide firm sales would be consistent with the findings of Hillberry and Hummels (2008), who document that U.S. manufacturers' shipments are extremely localized, with shipments within their 5-digit zip code of location being three times as large as shipments outside their zip code. Díaz-Lanchas et al. (2013) find evidence of an even stronger "own-zip-code" home bias using a micro-database of road freight shipments within Spain. Consistent with this prior literature, our first-stage results indicate that our instrument is indeed *relevant*, in the sense that the change

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[6]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 below, 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.

in the municipality-level stock of vehicles per capita between 2002-08 and 2009-13 has significant predictive power for the domestic (i.e., Spain-wide) sales of firms producing in that municipality.

Armed with these first-stage results, we 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, these IV estimates are significantly larger in absolute value than the OLS ones. This is consistent with the biases predicted by our baseline Melitz (2003)-type model in the plausible scenario in which our specification only imperfectly controls for a firm's supply and export demand determinants. Specifically, our IV estimates point at an intensive-margin elasticity of exports to domestic sales in the neighborhood of −1.6, while the OLS one is around −0.2.[7]

As indicated by condition (iii) above, a potential challenge to our identification approach is that the "boom-to-bust" changes in the stock of vehicles per capita in the municipality of location of a firm 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. Although, by definition, we cannot test this identification assumption, we provide several additional pieces of evidence that are consistent with the empirical relevance of the "vent-for-surplus" hypothesis and that address some specific sources of endogeneity that could affect the validity of the instrument in our baseline specification.

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. According to this hypothesis, if a significant share of vehicles is sold in the near 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 purchases of new vehicles. Our results are robust to this identification threat. Both the relevance of our instrument as well as the finding of a sizable negative elasticity between domestic sales and exports are robust to excluding from the estimating sample: (a) all firms in the auto industry, no matter where they are located; (b) all firms located in any zip code that hosts at least one auto-maker employing more than 20 workers; (c) all firms located in any zip code that is geographically close to a zip code in which a significant share of manufacturing employment is in the auto industry; and (d) all firms producing in sectors that are either leading input providers or leading buyers of the vehicles manufacturing industry.

Second, the "vent-for-surplus" hypothesis suggests that the elasticity of a firm's Spain-wide sales with respect to changes in local demand is likely to vary across firms in ways that can be verified. For instance, firms will naturally differ in their exposure to demand changes in their municipality of location depending on the share of their total domestic sales that is earned in that municipality. While we do not observe firms' sales distribution across different Spanish municipalities, it seems plausible that small firms will be more likely to concentrate their sales in their municipality of location than large firms. We indeed find that the first-stage elasticity of domestic sales with respect to our demand proxy is larger for smaller firms. We also find that a reduction in the municipality-level stock of vehicles per capita is associated with a larger reduction in Spain-wide sales for firms belonging to less "tradable" sectors, as measured by the sectoral share of within-

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[7]When estimating the effect of a *demand-driven* drop in domestic sales on the probability of exporting, we find an estimate that is not statistically different from zero.

province shipments in total shipments (computed from the C-Intereg database on road freight shipments within Spain).

Third, because different geographic areas in Spain were affected by the Great Recession in very heterogeneous degrees, it is conceivable that for many firms the "vent-for-surplus" mechanism would have operated largely at the *intranational* level. Rather than being pushed towards export markets, certain firms located in areas with disproportionate decreases in local demand could have redirected their sales largely towards other regions *within* Spain in which local demand decreased less (or increased). This implies that we should observe a larger elasticity of firms' Spain-wide sales with respect to proxies that capture changes in demand at the *province level* (rather than at the municipality level), as they preclude firms from redirecting their sales across municipalities belonging to the same province.[8] Conversely, if one were to hypothesize that our measure of changes in the stock of vehicles per capita is purely operating as a proxy for changes in unobserved marginal cost shifters (e.g., unobserved factor prices), then any dispersion in these unobserved shifters across municipalities located in the same province would imply that a firm's domestic sales elasticity with respect to our province-level instrument should be *smaller* than that with respect to our municipality-level instrument, as the province-level instrument would naturally be a worse proxy for the unobserved marginal costs shifters relevant to the firm. Our results in fact feature more than twice as large a response of domestic sales to a change in the instrument when the latter is measured at the province level than when it is measured at the municipality level.[9]

Fourth, consistently with the hypothesis that firms face increasing marginal costs of production and that the slope of these costs is inversely related to the elasticity of output with respect to inputs whose investment is not pre-determined or irreversible, we document that the estimated causal effect of demand-driven changes in domestic sales on exports is smaller for firms in labor-intensive and material-intensive sectors, suggesting the importance of fixed factors and capacity utilization in explaining this causal linkage.

Fifth, while our baseline instrumentation approach exploits a proxy for demand changes and is thus agnostic about the underlying causes of the differential impact of the Great Recession in Spain, we also explore alternative instrumentation strategies that focus instead on the deep roots of the differential fall in demand across Spanish regions. More specifically, we show that, relative to the boom years, firm-level domestic sales fell by more in municipalities with lower housing supply elasticities (in which house prices grew disproportionately during the boom years), in zip codes with a larger pre-crisis contribution of the construction sector to total labor income, and in provinces that experienced larger declines in tourism during the bust years.[10] Reassuringly, the second-stage elasticities of exports to domestic sales associated with these instruments are similar in magnitude to those obtained with our benchmark instrument.

Sixth, although we control for firm-specific average wages in all of our specifications, compositional changes in the firm's workforce may have caused changes in effective labor costs that our

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[8]While there are over 8,000 municipalities in Spain, there are only 50 provinces. Provinces are therefore significantly larger than municipalities.

[9]Consistently with changes in the stock of vehicles per capita capturing demand changes, we also find that a firm's domestic sales react to a distance- and population-weighted average of the changes in the stock of vehicles in all municipalities other than the municipality in which the firm is located, even after controlling for the changes in the stock of vehicles in the firm's municipality.

[10]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.

wage measure does not correctly capture. 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. We do indeed observe that firms whose share of temporary workers dropped by more in the bust relative to the boom experienced a smaller drop in their exports, consistently with the hypothesis that an increase in the ratio of permanent to temporary workers had an effect equivalent to a positive supply shock. The elasticity of exports with respect to domestic sales remains however 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 the firms does not change the second-stage estimate of the elasticity of exports with respect to domestic sales.

Seventh, and finally, we address the possibility that the correlation between boom to bust changes in the firm's domestic sales and in the stock of vehicles per capita is spurious and due to the presence of cross-municipality correlation in these variables' time trends. To rule out this possible explanation for our results, we perform a placebo exercise in which we break each of the boom and the bust periods into two subperiods, and evaluate whether our instrument (changes in demand between the boom and bust periods) predicts changes in domestic sales between the two boom subperiods (i.e., between 2006-08 and 2002-05) and between the two bust subperiods (i.e., between 2012-13 and 2009-11). In both cases, we find that it does not.

Having established a causal link between changes in domestic demand and exports that operates through firms' changes in domestic sales, 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 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, and employ the resulting estimates to quantitatively evaluate the importance of the "vent-for-surplus" mechanism in explaining the 2009-13 observed 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 54.7% 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*.[11] 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

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[11] A broader search to include top general-interest journals identified Neary and Schweinberger (1986), who provide a neoclassical rationale for the "vent-for-surplus" idea.

empirical evidence. This literature includes the work of Vannoorenberghe (2012), Blum et al. (2013), Soderbery (2014), and Ahn and McQuoid (2017). The results in those papers very much resonate with the OLS results using yearly data that we describe in Online Appendix G. Relative to this prior literature, our paper provides a more explicit discussion of the endogeneity concerns associated with simple OLS reduced-form regressions. More importantly, our paper also attempts to identify and structurally interpret the causal effect of a domestic slump on exporting by exploiting plausibly exogenous variation in domestic sales during a particularly salient episode. Relatedly, in contemporaneous work, Fan et al. (2018) 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* (reverse) 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. For these reasons, even if one takes their findings at face value, it would be unreasonable to interpret them as questioning the empirical relevance of the "vent-for-surplus" mechanism.[12]

Our identification strategy is inspired by the influential work of Mian and Sufi (and collaborators) on the causes and consequences of the Great Recession in the United States. Specifically, Mian and Sufi (2009) identify important differences in the extent to which the mortgage default crisis affected household wealth in different areas of the United States. In subsequent work, Mian, Rao and Sufi (2013) and Mian and Sufi (2014) study how the unequal geographic distribution of household wealth losses resulting from the housing crisis gave rise to a geographically unequal decline in consumption across U.S. counties. Our finding that geographical variation in the change in the stock of vehicles per capita is a significant predictor of variation in the change in local manufacturing sales in Spain is very much consistent with the findings in Mian and Sufi (2013), who also explore the link between household housing wealth and auto sales. Illustrating this link in the Spanish case would be interesting, but this is complicated by the sluggish adjustment of house prices in Spain during the financial crisis, as documented among others by Akin et al. (2014).[13]

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) and discuss its implications for the estimation of the causal impact of demand-driven changes in domestic sales on exports. In Section 3, we introduce our firm-level data and, in Section 4, we develop our core instrumental variable estimation approach. The results of this instrumental variable approach are presented in Section 5. We present additional evidence in favor of the "vent-for-surplus" mechanism in Section 6. In Section 7, we generalize the baseline model à la Melitz (2003) to allow for non-constant marginal costs, and use this framework to quantify the importance of the "vent-for-surplus" channel in linking the slump in domestic sales to the growth in Spanish exports. We offer some concluding remarks in Section 8.

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[12]Our paper also relates to previous work documenting 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). This literature is largely descriptive and has not attempted to test the relative contribution of different mechanisms in explaining the patterns observed in the data.

[13]More specifically, the fact that the housing market adjustment following the bursting of the housing bubble in Spain was largely made through quantities rather than prices implies that standard measures of housing wealth in Spain are not as good predictors for household consumption as they are in other countries.

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 and our choice of an adequate estimator, we first consider the implications for this question of a model of exporting with heterogeneous firms along the lines of Melitz (2003), which is 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 our evidence contradictory with this assumption, in Section 7 we will develop 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, firms' production locations within Spain by ℓ, 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. At a given point in time, 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 sectoral price index in j, E_{sj} is the total sectoral expenditure in market j expressed in units of the numeraire; 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 is a measure of firm-specific productivity, and ω_i is the firm-specific cost of a bundle of inputs.[14] 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:

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[14]Since our econometric specifications below include some location fixed effects, it would be straightforward to let the iceberg trade cost τ_{sj} also be a function of the production location ℓ .

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

The bulk of our empirical analysis will compare firm-level export behavior in a *bust* period, relative to a *boom* period.[15] 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

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

In order to transition to an estimating equation, we model the change in firm-specific foreign demand, productivity and cost levels as follows:

$$
\begin{aligned}
\Delta \ln(\xi_{ix}) &= \xi_{sx} + \xi_{\ell x} + u_{ix}^{\xi}, \\
\Delta \ln(\varphi_i) &= \varphi_s + \varphi_\ell + \delta_\varphi \Delta \ln(\varphi_i^*) + u_i^{\varphi}, \\
\Delta \ln(\omega_i) &= \omega_s + \omega_\ell + \delta_\omega \Delta \ln(\omega_i^*) + u_i^{\omega}.
\end{aligned}
\qquad (5)
$$

Note that we are decomposing these terms into (i) a sector fixed effect, (ii) a production location fixed effect, (iii) an *observable* part of these terms for the case of productivity (φ_i^*) and for input bundle costs (ω_i^*), and (iv) a residual term.[16] We can thus re-write equation (4) as:

$$\Delta \ln R_{ix} = \gamma_{sx} + \gamma_{\ell x} + (\sigma - 1)\delta_\varphi \Delta \ln(\varphi_i^*) - (\sigma - 1)\delta_\omega \Delta \ln(\omega_i^*) + \varepsilon_{ix}, \qquad (6)$$

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

$$\varepsilon_{ix} = (\sigma - 1)[u_{ix}^{\xi} + u_i^{\varphi} - u_i^{\omega}]. \qquad (7)$$

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

$$\Delta \ln R_{id} = \gamma_{sd} + \gamma_{\ell d} + (\sigma - 1)\delta_\varphi \Delta \ln(\varphi_i^*) - (\sigma - 1)\delta_\omega \Delta \ln(\omega_i^*) + \varepsilon_{id}, \qquad (8)$$

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

$$\varepsilon_{id} = (\sigma - 1)[u_{id}^{\xi} + u_i^{\varphi} - u_i^{\omega}]. \qquad (9)$$

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: changes to ξ_{id} that are independent of changes in the other model fundamentals (i.e. ξ_{ix}, φ_i, and ω_i) have no effect on $\ln R_{ix}$. However, many estimators of the impact of log domestic sales on log exports based on observational data will yield estimates that

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[15]In Online Appendix G, we theoretically develop and empirically test specifications that use yearly data; these results facilitate the comparison of our estimates with those in the previous literature.

[16]More precisely, we assume that $\Delta \ln \xi_{ix} + \Delta \ln \varphi_i - \Delta \ln \omega_i = d_s + d_\ell + \delta_\varphi \Delta \ln(\varphi_i^*) + \delta_\omega \Delta \ln(\omega_i^*) + u_i$, with u_i incorporating the unobserved components of export demand, productivity and factor costs, and $\mathbb{E}[u_i | \{d\}_s, \{d\}_\ell, \Delta \ln(\varphi_i^*), \Delta \ln(\omega_i^*)] = 0$, where $\{d\}_s$ denotes a complete set of sector-specific dummy variables, and $\{d\}_\ell$ is a complete set of location-specific dummy variables.

differ from zero, 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):

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

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

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

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 $\{d\}_s$, location fixed effects $\{d\}_{\ell}$, and the observable covariates $\Delta \ln \varphi_i^*$, and $\Delta \ln \omega_i^*$.

We draw two main conclusions from equation (11). First, as long as changes in productivity and production factor costs are not perfectly observable – and their unobserved component is not fully captured by the sector and location fixed effects – there will be a spurious 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 spurious positive correlation will lead $\hat{\beta}_{OLS}$ to be biased upwards. Second, even when 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. $cov(u_{ix}^{\xi}, u_{id}^{\xi}) \neq 0$), the OLS estimator of β will also converge to a non-zero value. Because this residual demand does not capture market-specific aggregate shocks (which are controlled by the sectoral fixed effects), it seems plausible that u_{ix}^{ξ} and u_{id}^{ξ} will be positively correlated in the data, leading $\hat{\beta}_{OLS}$ again to be biased upwards. Notice also that, if we had not controlled for sectoral and location fixed effects, the probability limit of the OLS estimator of β would likely be even larger.[17]

Consider next using an IV estimator of the parameters in equation (11). 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

$$plim(\hat{\beta}_{IV}) = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \mathcal{Z}_{id})}{cov(\Delta \ln \mathcal{R}_{id})} = \frac{cov(u_{ix}^{\xi} + u_i^{\varphi} - u_i^{\omega}, \mathcal{Z}_{id})}{cov(u_{id}^{\xi} + u_i^{\varphi} - u_i^{\omega}, \mathcal{Z}_{id})}, \qquad (13)$$

where, as above, we use $\mathcal{Z}_{id}$ to denote the residual from projecting Z_{id} on a set of sector and location fixed effects, and on the observable covariates $\Delta \ln \varphi_i^*$, and $\Delta \ln \omega_i^*$. The constant-marginal-cost

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[17]To give an example, the probability limit of $\hat{\beta}_{OLS}$ in the absence of production location fixed effects is:

$$plim(\hat{\beta}_{OLS}) = \frac{cov(u_{ix}^{\xi} + \xi_{\ell x} + \varphi_{\ell} - \omega_{\ell} + u_i^{\varphi} - u_i^{\omega}, u_{idt}^{\xi} + \xi_{\ell d} + \varphi_{\ell} - \omega_{\ell} + u_i^{\varphi} - u_i^{\omega})}{var(u_{idt}^{\xi} + \xi_{\ell d} + \varphi_{\ell} - \omega_{\ell} + u_i^{\varphi} - u_i^{\omega})}, \qquad (12)$$

which is likely larger than the expression in equation (11) due to: the presence of $\varphi_{\ell} - \omega_{\ell}$ in both terms of the covariance in the numerator of the expression in equation (12); and, the likely positive correlation between $\xi_{\ell x}$ and $\xi_{\ell d}$.

model predicts that $\hat{\beta}_{IV}$ converges in probability to its true value of zero as long as the instrument $\mathcal{Z}_{id}$ verifies two conditions: (a) it is correlated with the change in domestic sales of firm i after controlling for (or partialing out) sector and location 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 (13), an instrument can only (generically) verify conditions (a) and (b) if its effect on domestic sales works *exclusively* through the change in domestic demand not accounted for by the fixed effects and observable covariates included in the estimating equation, i.e., u_{id}^ξ.

Although our discussion above has centered around the role of unobserved supply and 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 total sales of firms, it is important to stress that measurement error in this setting does *not* just lead to attenuation bias as in the classical error-in-variables model. More specifically, and as we detail in Appendix A.1 (see also Berman et al., 2015), under plausible conditions, measurement error in firm total sales and exports will lead to a negative bias in the OLS estimate $\hat{\beta}_{OLS}$. As we show in the Appendix, however, if an instrument satisfies the same conditions (a) and (b) outlined above, and is also mean independent of the measurement error in exports, the IV estimator in equation (13) will still converge to zero in the presence of measurement error in total sales and exports.

We have focused our discussion so far on the intensive margin of exports, namely the impact of domestic demand shocks on the level of exports conditional on exporting. In Online Appendix E, 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 changes in domestic sales on the probability of exporting, even if the true effect were to be zero, one is likely to estimate a spurious positive elasticity whenever productivity and production factor costs are not perfectly captured by sector and location fixed effects and observable controls, or whenever unobserved *residual* demand shocks are positively correlated across 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 the various fixed effects and marginal cost proxies. Consequently, if the instrument affects domestic sales exclusively through the demand shock u_{id}^ξ, it will continue to be valid in those extensive margin specifications (see Online Appendix E for more details).

3 Setting and Data

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

Figure 2: The Great Recession in Spain
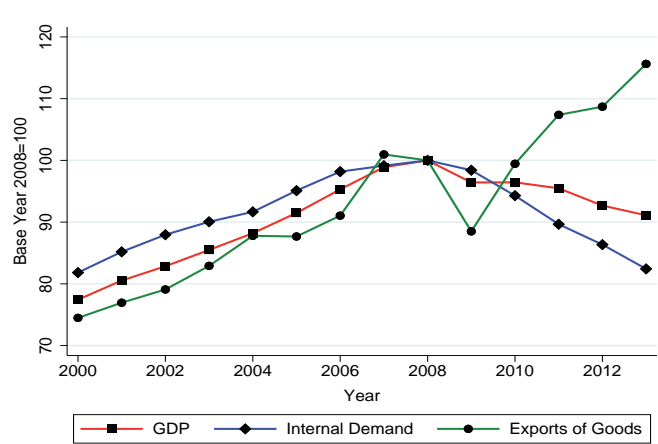
3.1 The Great Recession in Spain: Description

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 GDP and internal demand grew by approximately 20% in real terms.[18] 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%. In that same period, the unemployment rate shot up from 9% to 26%.

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.[19] For instance, in 2006, 735,000 new houses were built in Spain, a number comparable to that in Germany, Italy and the UK *combined*.[20] 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). The ratio of mortgage credit to GDP went up from 40% in 2000 to 100% in 2008 (Basco and Lopez-Rodriguez, 2018). Importantly, the very high loan-to-value (LTV) ratios associated with mortgage credit were partly used 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. The recession officially started in the fourth quarter of 2008, and intensified during 2009 with a 3.6% annual drop in 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. In fact, in 2013, the national stock of vehicles in Spain was *lower* than in 2008 by around 52,000 units.

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[18]Internal 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.

[19]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, with a similar pattern for the contribution of this sector to Spain's GDP (12.4% in 2007 and 6.8% in 2014).

[20]See EU Buildings Database at https://ec.europa.eu/energy/en/eu-buildings-database.

Importantly for the identification strategy we describe in the next section, the real estate boom and subsequent bust featured significant geographic variation, concentrating mainly in some parts of the Mediterranean coast and in medium-sized and large cities. As we shall document in the next section, this in turn translated into substantial geographic variation in the extent to which the Great Recession affected domestic demand and 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 2008-2013 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 2009-2013 at an even faster rate than during the boom years. Specifically, while exports had grown by an accumulated 34% in the eight-year period 2000-2008, 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 23.33% in 2013.

One might wonder whether changes in international relative prices could explain the growth in Spanish exports during the period 2009-2013. It is however easy to rule out exchange rate movements as a key operating mechanism since, as shown in Figure 1, Spanish exports clearly outperformed those of other countries in the euro area (even though Spain's GDP dropped faster than the euro area average). It has also been argued that Spain underwent an internal devaluation (through wage moderation starting in 2009, and via a labor market reform in 2012), but there is little evidence that export prices in Spanish manufacturing fell relative to export prices in other euro area countries in the period 2009-2013.[21]

Motivated by these facts, we will hereafter focus on an exploration of the "vent-for-surplus" mechanism, according to which the domestic slump, by freeing up production capacity, might have *directly* incentivized Spanish producers to sell their goods in foreign markets. In principle, the associated 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 foreign markets). Later in the paper, we will explore both margins, but 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 export growth between 2008 and 2013, while the remaining 86% was driven by continuing exporters. Similarly, in our sample of manufacturing firms, 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.[22]

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[21]More specifically, Eurostat data on unit values indicate a very small decline of 0.2% between 2008 and 2013 in Spanish export prices relative to those in the Euro area. Conversely, Eurostat data on industrial producer prices for non-domestic markets reveal an *increase* of 2.5% in Spanish relative export prices.

[22]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. See Section 3.3 for a description of our data limitations.

3.3 Data Sources

Our data cover the period 2000-2013 and come from two separate confidential administrative data sources. The first is the Commercial Registry (*Registro Mercantil Central*). It contains the annual financial statements of around 85% of registered firms in the non-financial market economy in Spain.[23] Among other variables, it includes information on the following: 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), and total fixed assets.[24]

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 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, 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 requires 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, information must be reported for all transactions performed by a firm during a natural year as long as at least one of these transactions exceeds 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 have at least one transaction exceeding 50,000 euros in that year (for more details see Appendix B).

In both datasets, a firm is defined as a business constituted in the form of a Corporation (*Sociedad Anónima*), a Limited Liability Company (*Sociedad Limitada*), or a Cooperative (*Cooperativa*). We merge both datasets using the fiscal identifier of each firm. Using the merged database, we define each firm's domestic sales as the difference between its total annual net operating revenue and its total export volume, which motivated our discussion of measurement error in Section 2.

To confirm the validity of the information contained in the resulting dataset, we compare its coverage with the official publicly available aggregate data on output, employment and total wage bill (from National Accounts) and on goods exports (from Customs). Figure 3 shows that our dataset tracks nearly perfectly the aggregate evolution over time of output, employment, total

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[23]We obtain information on the Commercial Registry from two different sources: (i) the *Central de Balances* dataset, compiled by the Bank of Spain, and (ii) the *Sabi* dataset, compiled by Informa (a private company). For details on how we combine these two datasets, see Almunia, Lopez-Rodriguez and Moral-Benito (2018).

[24]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 sector classification during the period of analysis. 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.
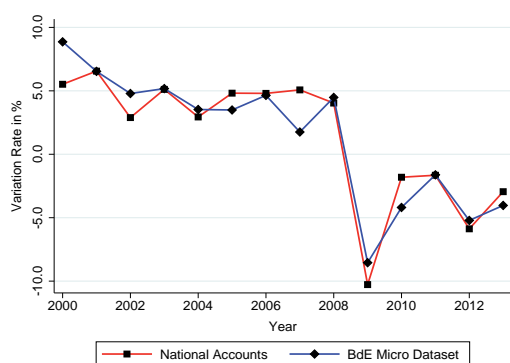
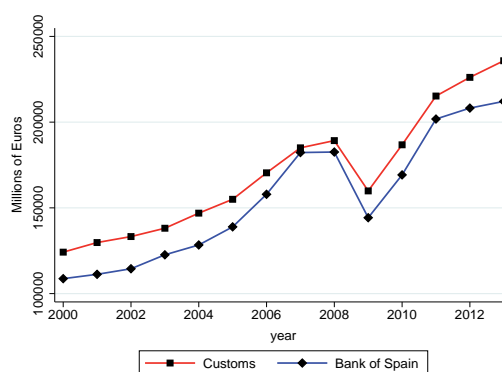Figure 3: Output, Employment, Wage Bill and Export Dynamics
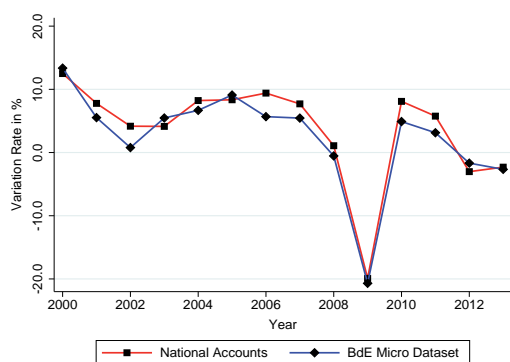
Panel (a): Output
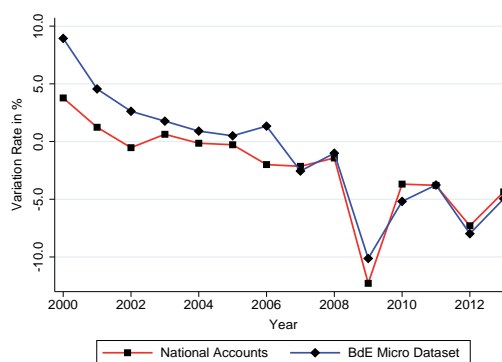
Panel (b): Employment

Panel (c): Wage Bill

Panel (d): Exports

payments to labor, and exports. Due to the reporting thresholds described above, aggregate exports in our sample naturally fall a bit short of aggregate exports in the Customs data, but note that the gap is very similar in the boom and bust periods (the average coverage is 91.8% in 2000-08 and 91.3% in 2009-13).[25]

We complement the firm-level data described above with yearly municipality-level data on the stock of vehicles and on total population. The information on the stock of vehicles by municipality is provided by the Spanish Registry of Motor Vehicles (*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. More precisely, 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 for the stock of vehicles and population to all firms located in the same municipality, independently of the zip code of location; for firms in zip codes containing multiple municipalities, we construct a zip code-level instrument by averaging the stock of vehicles per capita across these municipalities.

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[25]Most of the gap in coverage is explained by the fact that a nontrivial share of Spanish exports recorded by Customs 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 2010-2013 (own calculations based on public Customs data).

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 approach and how we seek to address them.

4.1 Identification Approach

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 4 illustrates this fact. The figure plots the standardized percentage change in domestic sales for the average firm located in each of the 47 Spanish peninsular provinces and operating in at least one year of the boom period (2002-2008) and at least one year of the bust period (2009-2013).[26] 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. Specifically, Figure 4 illustrates that firms located in the northern and western regions saw changes in domestic

Figure 4: The Great Recession in Spain: Variation Across Provinces
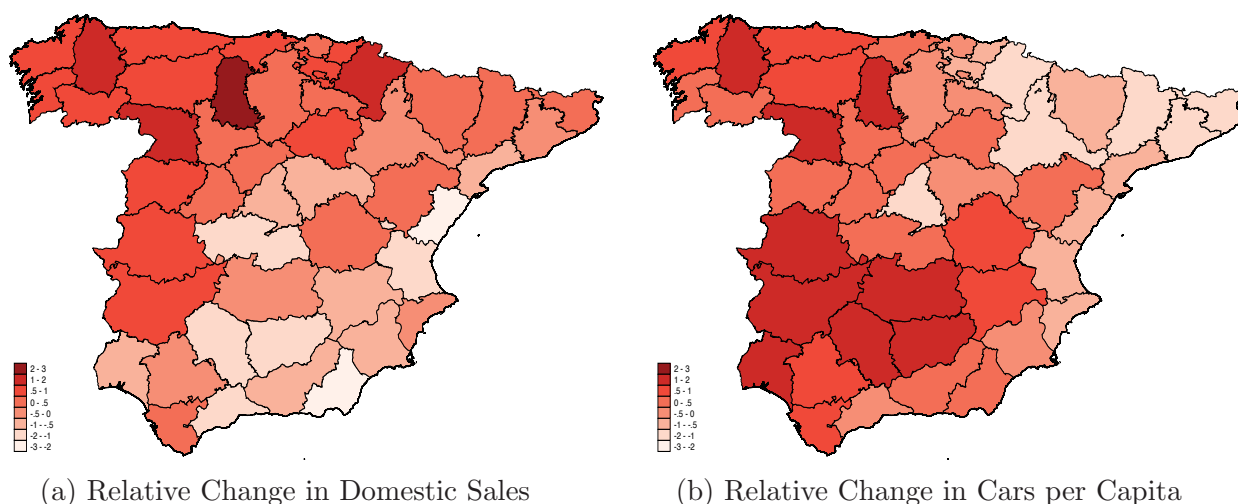
(a) Relative Change in Domestic Sales   (b) Relative Change in Cars per Capita

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 province, it means that the average firm located in this province 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 mean province. 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 province, it means that this province 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 mean province.

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.

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[26]Figure C.2 in Appendix C.2 shows the annual average number of firms and exporters by province for the period 2002-2008. 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).

The heterogeneity in the changes in domestic sales that we document in panel (a) of Figure 4 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 attempting to measure 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. Panel (b) in Figure 4 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.[27] 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 cost experienced a relative reduction.

By illustrating provincial averages, the maps in Figure 4 hide substantial spatial variation at the sub-province level (across 5-digit zip codes) in both the boom-to-bust changes in average firm-level domestic sales and in the boom-to-bust changes in the number of vehicles per capita. We illustrate this variation in Figure 5 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 5 with the across-province variation illustrated in Figure 4, the average zip code changes illustrated in Figure 5 have been standardized using the Spain-wide mean and cross-province standard deviation used to standardize the corresponding variables in Figure 4.

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 small reductions in firm average domestic sales (relative to the Spain-wide average), surrounding zip codes experienced changes in domestic sales 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 of Madrid saw increases in vehicles per capita significantly above the national average. Panels (c) and (d) provide analogous information for the region of Barcelona. Although the heterogeneity across zip codes located in the province of Barcelona 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.

In the next section, we exploit the variation illustrated in Figures 4 and 5 to identify the impact of a local demand shock 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 domestic sales in the bust period relative to the boom period is associated with a relative increase in export sales between these two periods. With this aim, we will use observed "boom-to-bust" changes in the stock of vehicles per capita at the zip code level as a proxy for the changes in the aggregate demand for

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[27]Changes in the number 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 contamination from the "cash for clunkers" program (*Plan PIVE*) that the Spanish government put in place during the bust period.

manufacturing goods that the corresponding geographical area experienced in the bust relative to the boom period. Equipped with this proxy for local goods demand, we will use it to instrument for *firm-level* changes in domestic sales for firms located in that zip code.

Our identification strategy is based on three main pillars. First, it builds on the fact that durable goods consumption, and vehicle purchases in particular, are strongly procyclical and thus are a useful proxy for changes in 'local demand', i.e., the overall propensity of an area's inhabitants to consume (see Stock and Watson, 1999). Consistent with this notion, Mian and Sufi (2013) document how 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 United States translated into geographical variation in vehicle purchases. It would be interesting to tie the geographical variation in the change in the stock of vehicles per capita in Spain to the housing slump, but idiosyncratic features of the Spanish housing market complicate such an analysis. In Section 6, we revisit this issue and explore the robustness of our results to an alternative shifter of firms' domestic sales that uses a determinant of the housing supply elasticity in a given zip code as a proxy for the magnitude of the negative impact of the Great Recession on household housing wealth and consumption.

The second building block of our identification strategy is that changes in zip code-level demand are a good predictor for changes in domestic (Spain-wide) sales of Spanish firms producing in the corresponding zip code. This would naturally be the case if domestic sales of firms were disproportionately localized in the zip code in which production takes place. Indeed, as mentioned in the Introduction, Hillberry and Hummels (2008) document the existence of such a 'zip code home bias' in U.S. manufacturers' shipments. Díaz-Lanchas et al. (2013) present analogous evidence for Spain. We cannot replicate these results with our data, but we document in Appendix C.3 the existence of significant home bias in manufacturing shipments with data at the *province* level.[28]

The third and final pillar of our identification approach is ensuring that changes in local vehicle purchases per capita are not correlated with supply shocks that might have an independent effect on the exporting decisions of Spanish firms. This exclusion restriction is central to the validity of our strategy, so we next outline how it might be violated and how we will deal with potential threats to identification.

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[28]More specifically, we find that own-province sales shares range from a low 18% in Transport Equipment (an industry we exclude from our analysis, as explained in the next section) to a high of 43% for Nonmetallic Minerals. The overall provincial home-bias in manufacturing is 28% (see Figure C.3 in Appendix C.3). The data on province-to-province shipments comes from the C-Intereg database (for details on this database, see Llano et al., 2010.)

Figure 5: The Great Recession in Madrid and Barcelona: Variation Across Zip Codes
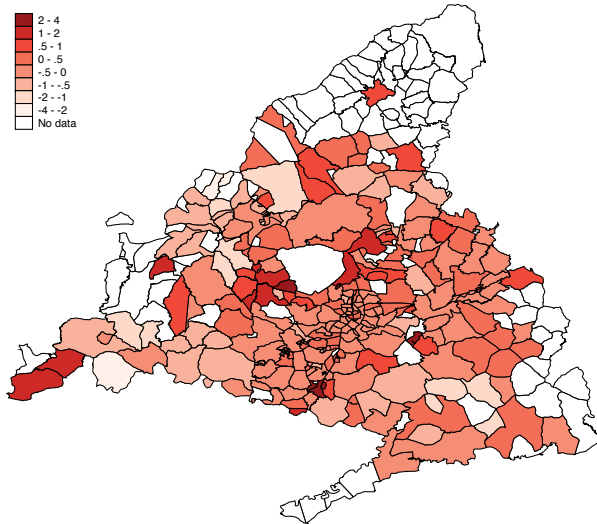
(a) Relative Change in Domestic Sales
(Madrid)
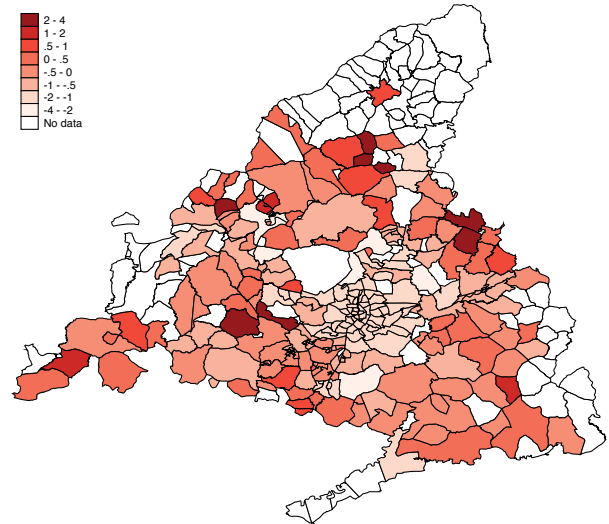
(b) Relative Change in Cars per Capita
(Madrid)

(c) Relative Change in Domestic Sales
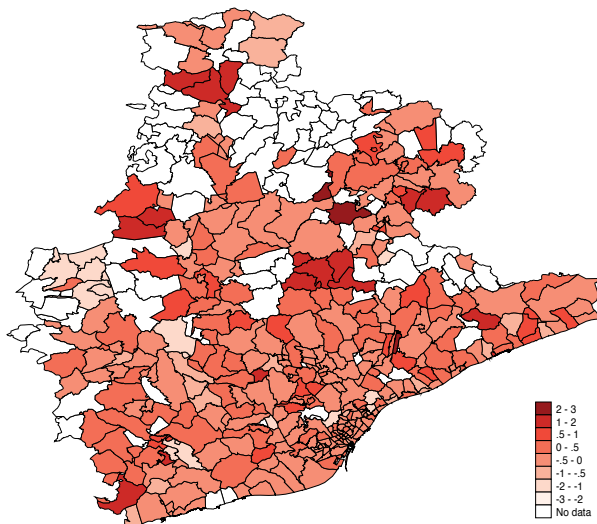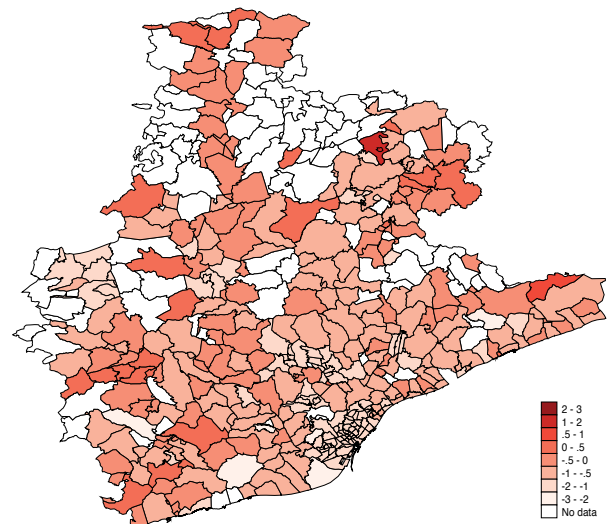(Barcelona)

(d) Relative Change in Cars 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".

4.2   Threats to Validity of the Instrument

The main concern with our approach is that the geographical variation in our demand measure might be correlated with geographical variation in unobserved supply shocks. While we cannot test this exclusion restriction formally, we address this endogeneity concern in two different ways.

First, we control in our specifications for sector and location (province) fixed effects and for firm-specific measures of productivity and labor costs. By controlling for sector fixed effects, 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 that experienced different changes in the stock of vehicles per capita. For example, these sector fixed effects control for shocks such as the expiration of the Multi Fiber Arrangement (MFA) on January 1, 2005, which eliminated all European Union quotas for textiles imported from China and which had a large impact on both the domestic sales and exports of Spanish textile manufacturers.[29] 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 regions in a manner that is correlated with our demand measure. This concern also motivates the inclusion of province fixed effects, with which we seek to control for unobserved variation in factor costs and productivity that is not picked up by our proxies for these variables.[30]

Our second approach to assuage endogeneity concerns is motivated by the fact that our various fixed effects and proxies for firm-level productivity and wage costs might not perfectly capture supply-side factors, and that unobserved, *residual* supply shocks might be correlated with our proxy for changes in local demand. For instance, if a disproportionate share of cars in Spain was sold in the municipalities in which car producing plants are located, then negative residual supply shocks affecting those car plants and their workers could well generate a correlation at the municipality level between car purchases and domestic sales. The Spanish motor vehicles sector represented on average around 7% of manufacturing employment during this period, so this is not an unreasonable concern, though we should stress that roughly 75% of cars purchased in Spain are imported (as indicated by data from the Spanish National Institute of Statistics). To deal with this threat to identification, in all regressions presented in the next section we exclude all firms operating in the auto industry (NACE Rev. 2 code 29). To further assuage this, in Section 6 we also explore 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 neighborhood 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.

In Section 6, we also perform several additional robustness tests of the vent-for-surplus hypothesis. Specifically, (i) we explore whether various heterogeneous effects are in line with what one would expect if the vent-for surplus mechanism was operating, (ii) we present IV estimates using alternative instruments for firms' domestic sales that exploit different variation in the data, (iii) we present regressions that control for several additional confounding factors, and finally, (iv) we perform falsification tests.

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[29]If sector fixed effects had not been included in our specifications and textile firms were to be on average located in Spanish regions that suffered larger negative local demand shocks, our estimates would confound the impact of the MFA expiration and the negative local shocks.

[30]As our instrument only varies at the municipality level and we use information only in one long time-difference, it is not feasible to introduce municipality fixed effects.

5 Baseline Results

5.1 Intensive Margin

Table 1 presents OLS estimates of the elasticity of "boom-to-bust" changes in firms' export flows 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 $8,018$ such firms in our dataset.

As discussed in Section 2, when no controls are included in the regression, we expect to observe a positive relationship between a firm's changes in domestic and foreign sales. This positive relationship is indeed observed in column 1 of Table 1, in which we estimate an elasticity of export flows with respect to domestic sales of 0.131. In the remaining columns of Table 1, we control for various sources of marginal cost heterogeneity across firms, with the aim of controlling for sources of correlation between firms' exports and domestic sales other than those captured by the vent-

Table 1: Intensive Margin: Ordinary Least Squares Estimates

Dependent Variable:			ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Domestic Sales)	0.131^a	-0.147^a	-0.228^a	-0.217^a	-0.204^a	-0.186^a
	(0.025)	(0.028)	(0.027)	(0.027)	(0.027)	(0.030)
ΔLn(TFP)		1.057^a	1.298^a	1.375^a	1.357^a	1.336^a
		(0.045)	(0.052)	(0.051)	(0.051)	(0.053)
ΔLn(Avg. Wages)			-0.590^a	-0.540^a	-0.525^a	-0.482^a
			(0.047)	(0.051)	(0.051)	(0.054)
Observations	8,018	8,018	8,018	8,018	8,018	7,507
R-squared	0.005	0.088	0.106	0.146	0.158	0.265
Sector FE	No	No	No	Yes	Yes	Yes
Province FE	No	No	No	No	Yes	No
Municipality FE	No	No	No	No	No	Yes

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, ΔLn(X) is the difference in Ln(X) between its average in the 2009-2013 period and its average in the 2002-2008 period. The estimation sample includes all firms selling in at least one year in the period 2002-2008 and in the period 2009-2013.

for-surplus mechanism. In columns 2 and 3, we control for log changes in firms' observed marginal costs. Specifically, in column 2 we control for the change in firms' productivity (estimated following the procedure in Gandhi et al., 2016, as detailed in Online Appendix F), and in column 3 we control for the change in firms' average wages (reported by the firm in its financial statement). 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 negative (-0.228), 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 are time varying. To do so, and motivated by the specification in equation (10), we sequentially add sector fixed effects (in column 4) and location fixed effects (in columns 5 and 6). In the latter case, we first include province fixed effects and, in column 6, we instead include municipality fixed effects. The resulting estimates continue to be negative and indicate that a 10%

decrease in a firm's domestic sales, keeping its productivity and average wages constant, implies around a 2% increase in its aggregate export flows.[31]

One might be concerned that, because firms' total sales are a key input in the computation of our TFP measure, our empirical results are just unveiling a mechanical negative correlation between exports and domestic sales once one holds total sales revenue constant (by controlling

Table 2: Intensive Margin: Two-Stage Least Squares Estimates

Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔLn(Domestic Sales)					-2.185[a]	-1.346[a]	-1.393[a]	-1.602[a]
					(0.622)	(0.359)	(0.350)	(0.437)
ΔLn(Vehicles p.c. in Municipality)	0.336[a]	0.447[a]	0.437[a]	0.363[a]				
	(0.060)	(0.066)	(0.065)	(0.068)				
ΔLn(TFP)		0.785[a]	0.948[a]	0.936[a]		1.991[a]	2.396[a]	2.657[a]
		(0.025)	(0.029)	(0.030)		(0.286)	(0.339)	(0.409)
ΔLn(Avg. Wages)			-0.544[a]	-0.447[a]			-1.225[a]	-1.149[a]
			(0.031)	(0.036)			(0.205)	(0.206)
F-statistic	31.17	46.39	45.00	28.32				
Observations	8,018	8,018	8,018	8,018	8,018	8,018	8,018	8,018
Sector FE	No	No	No	Yes	No	No	No	Yes
Province 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 municipality 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. Vehicles p.c. denotes the stock of vehicles per capita. 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(Vehicles p.c. in municipality) equals zero.

for it). Although log TFP and log total sales are obviously positively correlated (as one would expect in light of our model), the correlation is far from perfect, particularly when considering log changes in these variables. More specifically, the correlation between log changes in TFP and log changes in total sales in our yearly data is 0.31, while it is 0.54 when looking at boom-to-bust "long differences" in these variables. To further assuage this concern, in Section 6.4 we explore the robustness of our results to alternative measures of log firm TFP that feature a lower correlation with log firm sales.

In Table 2, we turn to our two-stage least squares 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. The first-stage estimates (reported in columns 1 to 4 and illustrated in Panel (a) of Figure C.4 in Appendix C.4) reveal that firms located in municipalities that experienced a larger drop in the stock of vehicles per capita 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 and for sector and province fixed effects: the statistic of an F-test for the null hypothesis

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[31]In Online Appendix G, we present OLS regressions using the full firm-year data for the period 2002-2013. Our results are quite similar to those in Table 1. Without controlling for supply factors, changes in domestic sales are positively associated with changes in exports. However, once we control for observable determinants of firms' marginal costs and for various fixed effects, we estimate a negative elasticity of exports to domestic sales. This elasticity is around -0.3 and thus somewhat larger (in absolute value) than the one obtained in our "long differences" specification. In Online Appendix G, we also explore variation in the elasticity of exports with respect to domestic sales across sectors.

that changes in the stock of vehicles per capita in a region have no impact on the domestic sales of the firms located in that municipality is comfortably above widely accepted critical values in all specifications.

The second stage estimates (reported in columns 5 to 8) reveal elasticities of exports with respect to domestic sales that are significantly larger (in absolute value) than the OLS elasticities reported in Table 1.[32] This is true regardless of whether one controls for sector and province fixed effects as well as for changes in our measures of the firm's productivity and labor costs. Our preferred estimate in column 8 indicates an elasticity of exports with respect to domestic sales of around -1.6. These significantly more negative IV elasticities are consistent with the hypothesis, formalized in equation (11), that, even after controlling for sector and location fixed effects and for firm proxies of productivity and average labor costs, there still remains unobserved determinants of firms' marginal costs that induce a spurious positive correlation between their sales in the domestic and foreign markets.

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 $\chi\%$, 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.[33]

5.2   Extensive Margin

We next turn to studying the causal impact of demand shocks on the extensive margin of exporting. As in our intensive margin regressions, we divide the sample period into a boom (2002-08) and a bust period (2009-13), and explore how demand-driven changes in domestic sales affect firms' probability of exporting in each of these two periods. More specifically, we implement a two-stage least squares estimator 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 and sector-period fixed effects, province-period fixed effects, the log of firm-level average TFP in that period, the log of average wages in that period, and the log of average domestic sales in that period, with log domestic sales instrumented with the average stock of vehicles per capita in the firm's municipality of location during that period.[34] Besides this linear probability model, we also estimate analogous specifications in which we substitute the dependent variable by a variable capturing the proportion of years in a given period (boom or bust) for which a firm exports.

The results are presented in Table 3. Column 1 reports the first stage for our full sample of 62,904 firms. As in Table 2, and given the inclusion of firm fixed effects, the results indicate that domestic sales fell more for firms located in municipalities with a larger decline in the stock of vehicles per capita. The F-stat (14.49) is, as in our intensive margin specifications, above standard critical values. Columns 2 and 3 then present OLS and IV estimates of the link between domestic sales and export status, while columns 4 and 5 report OLS and IV estimates of the link between

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[32]We illustrate the reduced-form relationship between the change in the log number of vehicles per capita of a municipality and the change in log exports of the firms located in the corresponding municipality in Panel (b) of Figure C.4 in Appendix C.4.

[33]The median export share among the 8,018 firms exporting in both boom and bust periods is 16.5%.

[34]Our results to an alternative specification in which the left-hand-side variable is a dummy variable that treats a firm as an 'exporter' only if it exports for two or more years in a given period.

domestic sales and the proportion of years exported. The results of these two specifications deliver very similar results. First, the OLS estimates in columns 2 and 4 suggest a positive relationship between domestic sales and the propensity to export. When isolating demand-driven variation

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

Dependent Variable:		Export Dummy		Proportion of Years	
	1st Stage (1)	OLS (2)	2nd Stage (3)	OLS (4)	2nd Stage (5)
Ln(Domestic Sales)		0.040^a (0.003)	-0.107 (0.181)	0.021^a (0.002)	-0.071 (0.094)
Ln(Vehicles p.c. in Municipality)	0.089^a (0.023)				
Ln(TFP)	1.075^a (0.016)	0.038^a (0.005)	0.196 (0.195)	0.050^a (0.003)	0.148 (0.101)
Ln(Average Wages)	-0.408^a (0.011)	-0.024^a (0.004)	-0.084 (0.074)	-0.031^a (0.003)	-0.068^c (0.038)
F-statistic	14.49				
Observations	125,808	125,808	125,808	125,808	125,808
Mean of Dep. Var.		0.183	0.183	0.113	0.113
Ext-Margin Elasticity		0.221	-0.584	0.181	-0.622

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by zip code appear in parenthesis. For any X, $\Delta Ln(X)$ 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. All specifications include firm fixed effects, sector-period fixed effects, and province-period fixed effects.

in domestic sales, however, the coefficient in column 3 turns negative and suggests that a 10% drop in domestic sales leads to a 1.07% increase in the probability of exporting. This effect is, however, very imprecisely estimated and it is thus not possible to reject the null hypothesis that demand shocks have *no* impact on the extensive margin of exporting. The same conclusion applies to column 5, which presents an estimate of the causal effect of demand shocks on the proportion of years exported.

Taken together, the results in Table 3 lead us to conclude that the vent-for-surplus mechanism did not appear to operate via the extensive margin. This result is perhaps not entirely surprising in light of the fact, discussed in Section 3.2, that more than 90% of the growth in Spanish exports during the bust period was explained by the intensive margin.

6 Robustness

In this section, we complement our baseline results with a series of robustness tests that further support the empirical relevance of the "vent-for-surplus" mechanism, and that address some specific sources of endogeneity that could affect the validity of our baseline identification strategy. Given the non-significant results obtained in Table 3 regarding the extensive margin of exports, we focus throughout this section on exploring the robustness of the intensive margin results in Table 2.

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

Dependent Variable:	Panel A: Exclude zipcodes w/ high auto employment share			Panel B: Exclude zipcodes with at least one sizeable auto maker		
	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	1st Stage	2nd Stage	OLS	1st Stage	2nd Stage
ΔLn(Domestic Sales)	-0.218^a		-2.382^a	-0.235^a		-2.787^a
	(0.030)		(0.535)	(0.038)		(0.694)
ΔLn(Vehicles p.c. in Municipality)		0.328^a			0.318^a	
		(0.075)			(0.088)	
ΔLn(TFP)	1.349^a	0.936^a	3.361^a	1.348^a	0.898^a	3.624^a
	(0.057)	(0.032)	(0.501)	(0.072)	(0.043)	(0.630)
ΔLn(Avg. Wages)	-0.487^a	-0.436^a	-1.428^a	-0.526^a	-0.408^a	-1.561^a
	(0.052)	(0.037)	(0.253)	(0.071)	(0.047)	(0.322)
F-statistic		19.04			13.11	
Observations	7,178	7,178	7,178	4,613	4,613	4,613

Dependent Variable:	Panel C: Exclude zipcodes 'neighboring' zipcodes in Panel A			Panel D: Exclude sectors w/ I-O links to automakers		
	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	1st Stage	2nd Stage	OLS	1st Stage	2nd Stage
ΔLn(Domestic Sales)	-0.204^a		-2.487^a	-0.179^a		-2.443^a
	(0.033)		(0.632)	(0.031)		(0.579)
ΔLn(Vehicles p.c. in Municipality)		0.332^a			0.339^a	
		(0.085)			(0.067)	
ΔLn(TFP)	1.324^a	0.928^a	3.428^a	1.298^a	0.933^a	3.396^a
	(0.062)	(0.032)	(0.583)	(0.063)	(0.035)	(0.538)
ΔLn(Avg. Wages)	-0.471^a	-0.404^a	-1.390^a	-0.520^a	-0.413^a	-1.448^a
	(0.058)	(0.037)	(0.276)	(0.060)	(0.037)	(0.252)
F-statistic		15.42			25.32	
Observations	6,137	6,137	6,137	6,080	6,080	6,080

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include sector and province fixed effects. For any X, $\Delta Ln(X)$ is the log difference between the average of X in 2009-2013 and its average in 2002-2008. 'Exp' denotes exports, and 'DSales' denotes domestic sales. 'Vehicles p.c.' denotes the stock of vehicles per capita. 'F-statistic' denotes the corresponding statistic for the null hypothesis that the coefficient on the ΔLn(Vehicles p.c. in Municipality) covariate is equal to zero.

6.1 Further Purges of the Auto Industry

While the sample used to compute the estimates in Table 2 excludes firms classified in the manufacturing of motor vehicles sector (see Section 4.2), one might still be concerned that the salient presence of firms in 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 in the unobserved *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, reducing the aggregate labor demand in this municipality.[35] Notice however that this source of endogeneity in our instrument would cause our baseline two-stage least squares estimates presented in Table 2 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 4 our two-stage least squares estimates 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 'neighboring' zip codes sharing the first four digits with a zip code that ranks in the top 25% of zip codes by share of manufacturing employment in motor-vehicles producers. Finally, in Panel D, we exclude all firms 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. The results in all panels point at slightly larger estimated elasticities (in absolute value), consistently with the notion that these sample restrictions attenuate concerns about our estimates being up upward biased. Notice however that, as a consequence of the reduction in sample sizes, the standard errors also increase significantly.

6.2 Heterogeneous Effects

Our first-stage and second-stage intensive margin specifications in Table 2 do not allow for heterogeneity in how firms are affected by local demand shocks and how they might respond to them. However, the 'vent-for-surplus' interpretation of the results suggests plausible sources of heterogeneity in both the first-stage coefficient on the instrument and in the second-stage coefficient on the log change in domestic sales.

For instance, it seems reasonable to expect smaller firms to see their domestic (Spain-wide) sales being more impacted by municipality-specific demand shocks than larger firms. This means that, as long as it is the case that our instrument is truly capturing changes in local demand between the boom and bust periods, we should expect the elasticity of a firm's domestic sales with respect to our instrument to be larger for smaller firms.[36] This is indeed what we observe in columns 1 to 5 of Table 5. The elasticity of the boom-to-bust change in log domestic sales with respect to the log change in the stock of vehicles per capita in the municipality of location of the firm is around 0.5 for firms with less than 25 employees and around 0.3 when we consider only firms with more than 25 employees (columns 1 and 2 of Table 5). Restricting the sample further to firms with more than 50 employees does not have an statistically significant effect the elasticity of interest (column 3), but this one becomes close to 0.2 when we focus only on firms with more than 100 employees (column 4) and below 0.1 when we do so with firms with more than 200 employees (column 5). In the latest case, we actually cannot reject the null hypothesis that the first-stage elasticity of interest is zero.

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[35]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, which may be other firms located in the same municipality but operating in different industries.

[36]Notice that there would be no immediate reason to expect that one should observe this heterogeneity if it were to be the case that the boom-to-bust log change in the number of vehicles per capita in a municipality is exclusively operating as a proxy for unobserved supply shocks affecting the firms located in such municipality.

Table 5: Heterogeneous Effects: First Stage

| Sample: | Number of workers is in the interval: | | | | | Low Home Bias | High Home Bias |
| | $(0, 25]$ | $[26, \infty)$ | $[51, \infty)$ | $[101, \infty)$ | $[201, \infty)$ | | |
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1st-Stage Coefficient	0.516^a	0.316^a	0.331^a	0.223^c	0.074	0.305^a	0.437^a
	(0.117)	(0.073)	(0.095)	(0.121)	(0.154)	(0.094)	(0.085)
F-Statistic	19.47	18.78	12.19	3.37	0.23	10.60	26.57
Observations	2,641	5,376	3,190	1,672	790	4,768	3,249

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include sector and province fixed effects. Vehicles p.c. denotes the stock of vehicles per capita. First-stage coefficient and F-statistic refer to the elasticity of the change in the firm's log domestic sales with respect to the change in log vehicles p.c. in the municipality of location of the firm. For columns 1 to 5, the firm's number of workers is measured as the average across all years the firm appears in the sample. In columns 6 and 7, we classify firms into low and high home bias firms depending on whether the provincial home bias of their sector is below or above the sectoral median.

Whenever a municipality experiences a drop in demand, the 'vent-for-surplus' mechanism predicts that firms located in it will try to recoup the lost sales in some other market. However, the larger the trade costs of shipping goods from such municipality to other markets that a firm faces, the harder it is for this firm to shift their sales towards new markets and, thus, the larger the elasticity of their domestic (Spain-wide) sales with respect to changes in demand in their municipality of location is expected to be. We cannot measure trade costs for each firm but we use the share of total shipments of a sector that remain within the same province of the municipality of origin (i.e. provincial 'home bias') to proxy for those 'outward' trade costs at the sectoral level. We then classify firms into low and high home bias firms depending on whether the provincial home bias of their sector is below or above the sectoral median. Consistently with the prediction of the 'vent-for-surplus' mechanism, we find that the elasticity of the boom-to-bust change in domestic sales with respect to the change in the stock of vehicles per capita is larger (0.436 vs. 0.305) in those sectors that are more *inward-oriented* (see columns 6 and 7 in Table 5).[37]

Table 6 presents patterns of heterogeneity in the second-stage elasticity of the boom-to-bust change in exports with respect to the boom-to-bust change in domestic sales that are informative about the economic mechanisms underlying the relationship between demand-driven changes in domestic sales and exports. First, while our baseline specification in equation (10) imposes a constant elasticity between changes in exports and demand-driven changes in domestic sales, it is natural to expect the percentage change in exports following a given percentage change in domestic sales to be larger for firms with initially lower export shares.[38] As the results in columns 1 and 2 of Table 6 show, the elasticity of changes in exports to demand-driven changes in domestic sales is indeed larger for firms with a lower (pre-sample) propensity to export (elasticity of -3.034 versus -0.839 for firms with high export propensity).

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[37]Analogously to the discussion in footnote 36, the heterogeneity in the first-stage elasticities documented in columns 6 and 7 of Table 5 is hard to rationalize under a hypothetical interpretation of our baseline results that maintains that our instrument is operating as a proxy for unobserved residual supply shocks in a municipality.

[38]E.g. if a firm is attempting to recoup €100,000 in lost domestic sales, which constitute a 10% drop in domestic sales, the required percentage increase in exports will be larger if the firm initially exported €111,000 worth of goods (i.e., an initial trade share of 10%) than if it initially exported €250,000 (i.e., an initial trade share of 20%).

Table 6: Heterogeneous Effects: Second Stage

Sample	Low prov-sec Exp. share (1)	High prov-sec Exp. share (2)	Low Labor Elasticity (3)	High Labor Elasticity (4)	Low Materials Elasticity (5)	High Materials Elasticity (6)
2nd-Stage Coefficient	-3.034[b] (1.177)	-0.839[c] (0.460)	-1.606[b] (0.684)	-1.350[b] (0.645)	-2.019[b] (0.853)	-1.078[b] (0.523)
1st-Stage F-Stat.	10.23	24.98	19.76	12.22	9.90	20.55
Observations	4,005	4,009	3,914	3,914	4,100	3,711

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include sector and province fixed effects. Vehicles p.c. denotes the stock of vehicles per capita. First-stage coefficient and F-statistic refer to the elasticity of the change in the firm's log domestic sales with respect to the change in log vehicles p.c. in the municipality of location of the firm. In columns 1 and 2, we classify firms on the basis of province- and sector-specific export shares. In columns 3 to 6, we classify firms on the basis of firm-specific labor and materials elasticities computed following the procedure in Bilir and Morales (2018).

If the "vent-for-surplus" mechanism is important in explaining the growth in Spanish exports during the Great Recession, then one would 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.[39] Columns 3 and 4 test this hypothesis when we identify our flexible input as labor, and columns 5 and 6 do so when we consider materials to be our flexible input. Our results generally confirm that the elasticity of the change in exports to changes in domestic sales is indeed higher for firms having lower output elasticities with respect flexible inputs. One should notice however that, while the difference in the estimates is large when we classify firms according to their materials output elasticity (-2.019 vs. -1.078), it is much smaller when we do so according to their labor output elasticity (-1.601 vs. -1.350). This may reflect the rigidity of the Spanish labor market and the consequent difficulties that Spanish firms faced during the bust period to adjust downwards their labor force in reaction to the drop in domestic demand.

6.3 Alternative Instruments

We next revert back to the baseline specification discussed in Section 5.1, but explore the robustness of the results presented in Table 2 to alternative instruments.

We first construct two alternative instrumental variables that are analogous to our baseline one except for the fact that they measure the change in the stock of vehicles per capita not only in the municipality of location of a firm but also in surrounding municipalities. The first alternative instrument measures the change in the stock of vehicles per capita at the province level.[40] A change in demand in the province of location of a firm will directly affect a larger share of the firm's domestic sales than a change in demand affecting only the municipality of location; thus, as long as the change in the stock of vehicles per capita is actually a demand proxy, we should expect the elasticity of a firm's domestic sales with respect to our province-level instrument to be larger

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[39]See Appendix A.2 for a formalization of the link between the slope of the short-run marginal cost function and the elasticity of output with respect to flexible inputs.

[40]Provinces are significantly larger than municipalities: while there are over 8,000 municipalities in Spain, there are only 50 provinces.

than our baseline elasticity with respect to our municipal-level instrument. This is indeed what we find when we compare the first-stage estimate in column 1 of Table 7 to that in column 4 of Table 2 (0.853 vs. 0.363).[41]

The second alternative instrument is a distance- and population-weighted sum of the change in the stock of vehicles per capita in all zip codes *other than* the zip code in which the firm is located.[42] If the change the stock of vehicles per capita was just a proxy for unobserved supply shifters, one would expect the change in vehicles per capita in a municipality to be a sufficient statistic for all relevant shifters affecting the firms located in such municipality. Reassuringly, the distance- and population-weighted sum of the change in the stock of vehicles per capita in other municipalities is correlated with firms' change in domestic sales even after controlling for the change in vehicles per capita in its own municipality (see column 3 in Table 7).

The different instrumental variables exploited so far rely on the change in the stock of vehicles per capita in a municipality being a *proxy* for demand changes in that municipality, but do not take a stance on the primitive sources or causes of the demand changes. We next construct alternative instruments that attempt to better capture the deep roots of the Great Recession in Spain. As described in Section 3.1, the Great Recession in Spain is largely driven by a real state bubble. Our third 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 1996 (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.[43] 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).[44]

Relatedly, we also use a fourth alternative instrument related to the construction sector. The burst of real state bubble affected directly the construction sector. As mentioned in footnote 19, 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 larger share of the workers employed in the construction sector during the boom ended up unemployed during the bust period. These workers saw their consumption capacity severely reduced in the bust period relative to the boom. Consequently, one may conjecture that the boom-to-bust drop in demand for manufacturing products was larger in those municipalities for which the construction sector was a particularly important

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[41]Conversely, if the correlation between firms' domestic sales and the change in the stock of vehicles per capita that we document in columns 1 to 4 of Table 2 was due to our instrument operating as a proxy for firms' unobserved marginal cost shifters, we would expect the elasticity of domestic sales with respect the province-level measure of our instrument to be lower than that with respect to the municipality-level measure, as the former would be a more noisy proxy of the supply factors relevant to the firm.

[42]In our baseline measure, we weight each zip code by its population divided by the logarithm of its distance to the zip code where the corresponding firm is located. We have experimented with alternative weights, such as dividing population by power functions of distances, and have found qualitatively very similar results.

[43]Indeed, we show in Appendix C.5 that there is a negative cross-sectional correlation between these housing supply elasticities and housing price growth during the boom years 2004-07.

[44]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.

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.[45]

Our fifth and last 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 in this case the number of foreign tourists that visited a province in 2002 divided by the population of the province in the same year.[46]

In Table 7, we report the results obtained under these different alternative instruments. As column 1 demonstrates, despite the differences in the first-stage coefficients on the province-level and the municipality-level measures of the change in the stock of vehicles per capita, our second-stage province-level results generate a response of exports to a fall in domestic (Spain-wide) sales very similar to that in our baseline (-1.425 vs. -1.602). Consequently, the Sargan test of overidentifying restrictions clearly fails to reject the null hypothesis that our instruments are valid. The results we obtain under the distance- and population-weighted vehicles per capita instrument in columns 2 and 3 are equally reassuring.

One might argue that the fact that we fail to reject the null hypothesis in the test of overidentifying restrictions performed in columns 1, 2 and 3 of Table 7 is not surprising, since these alternative instruments use a source of variation that is quite similar to that used in our baseline specification. In that respect, the results we obtain when using as instruments our observed housing supply determinant, the 2002 municipality-level construction wage bill or the 2002 province-level tourism share are more revealing. 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 sign of all first-stage coefficients is as expected, and that the second-stage elasticities of exports to domestic sales are quite similar in value (-1.595, -1.568, and -1.179, respectively) to those obtained with our benchmark instrumentation strategy in Table 2. Furthermore, the p-values of the Sargan test of overidentifying restrictions are very large (0.51, 0.99 and 0.97, respectively). 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.

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[45]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 shift-share 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 then that identification must come then from assumptions imposed on the distribution of the 2002 construction wage bill. See Goldsmith-Pinkham et al. (2018) for a discussion of identification in this context. Conversely, neither the identification approach in Borusyak et al. (2018) nor the discussion on inference in Adão et al. (2018) are applicable to our context.

[46]Considerations analogous to those in footnote 45 apply here.

Table 7: Alternative Instruments and Overidentification Tests

Dependent Variable:	ΔLn(Domestic Sales)					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Vehicles p.c. in Province)	0.853^a (0.223)					
ΔLn(Distance-Population Weighted Vehicles p.c. in Other Zip Codes)		0.260^a (0.028)	0.184^a (0.040)			
ΔLn(Vehicles p.c. in Municipality)			0.296^a (0.086)			
Ln(Urban Land Supply Ratio in 1996)				0.029^b (0.012)		
ΔLn(Construction Wage Bill) × 2002 Wage Bill Share in Municipality					0.331^a (0.054)	
ΔLn(Foreign Tourists) × 2002 Foreign Tourists p.c. in Province						0.280^a (0.098)
F-statistic	14.61	86.02	43.02	6.36	38.33	8.18

	ΔLn(Exports)					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Domestic Sales)	-1.425^a (0.400)	-1.628^a (0.527)	-1.336^a (0.395)	-1.595^c (0.927)	-1.568^a (0.535)	-1.179^a (0.257)
P-value for Sargan Test	0.46	0.34	0.34	0.51	0.99	0.97
Observations	8,018	7,949	7,949	6,940	7,928	8,018

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by province except for columns 4 and 5, in which they are clustered by municipality. 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, and columns 4 and 5 also include province fixed effects.

6.4 Controlling for Additional Confounding Factors

In spite of the controls included in our baseline specification, one may still be concerned that this specification might not be accounting for the effect of marginal cost shifters that could be correlated with our instrument, thus biasing our estimates. More specifically, one might be concerned that our firm-level measures of average wages and TFP are too crude to fully capture changes in firm-level supply conditions even when additionally controlling for sector and province fixed effects.

For instance, the dual nature of the Spanish labor market, with large differences in pay and job security between temporary- and permanent-contract workers, might have led certain firms to shed a disproportionate number of temporary, lower-paid workers during the bust. If so, conditional on the observed changes in our measure of average wages, our TFP measure could significantly underestimate the export potential of firms undergoing such skill-upgrading (or, at least, experience-upgrading). Similarly, our baseline specifications do not include any proxies for factor costs other than labor costs, yet it is likely that financial costs faced by firms (explicit via interest rates, or implicit via rationing) were also significantly impacted by the Great Recession.

Columns 1 to 4 of Table 8 present variants of our baseline specification that include controls for various alternative confounding factors related to the labor costs. Although all regressions reported

in Table 8 also include the controls included in our baseline specification, to save space, we do not include these estimates. To facilitate the comparison, we replicate in column 1 our baseline results in column 8 of Table 2. In column 2, we additionally control for the firm-level change in the share of temporary workers. The results suggest 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. The IV estimate of the causal effect of demand shocks on exporting is however only slightly lowered (elasticity of -1.443).[47] In columns 3 and 4, we introduce municipality-level controls for local labor market conditions. Column 3 includes the same change in

Table 8: Confounding Factors

Dependent Variable:	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Domestic Sales)	-1.602^a	-1.443^a	-1.655^a	-1.677^a	-1.416^a	-1.383^a	-1.667^a
	(0.437)	(0.434)	(0.480)	(0.478)	(0.450)	(0.435)	(0.481)
ΔShare of Temp. Workers (firm level)		-0.302^b					
		(0.118)					
ΔShare of Temp. Workers (municipality level)			-0.068				
			(0.194)				
ΔManufacturing Empl. p.c. (municipality level)				-0.266^a			
				(0.057)			
ΔLn(Financial Costs) (firm level)					-0.031^b		
					(0.014)		
Ln(Financial Costs in Boom) (firm level)						-0.000	
						(0.016)	
Ln(Financial Costs in Boom) (municipality level)							0.023
							(0.035)
F-Statistic	28.32	25.83	25.34	24.82	24.07	24.81	25.26
Observations	8,018	7,649	7,746	7,748	6,886	6,952	7,743

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

the ratio of temporary workers over total employment as in column 2, but computed with aggregate data 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 of these municipality-level variables has a significant effect on exporting.[48]

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 firms 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). As the results in column

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[47]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 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 causal impact of domestic shocks on exports is largely unaffected.

[48]More specifically, firms located in municipalities with larger declines in manufacturing employment per capita experienced higher export growth, presumably due to workers extra effort in reaction to the reduction in employment opportunities in their municipality.

5 illustrate, the impact of this measure on firms' changes in exports is not statistically different from zero, and including this variable has only a effect on the estimate of the elasticity of exports to domestic sales (which drops to -1.416). 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. No matter whether we measure financial costs in the boom using each firm's information (in column 6) or as the average financial costs of all other firms located in the same municipality (in 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 financial costs in the boom has little empirical support.

6.5 Alternative Productivity Estimates

We next test the robustness of our results to alternative approaches to measuring firms' productivity. Columns 1 and 2 in Table 9 replicate our baseline OLS and IV estimates presented in column 6 of Table 1 and column 8 of Table 2, respectively, and 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, both productivity measures exploit the assumptions 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. In our baseline approach, we assume that the firm's production function is a Leontief aggregator of materials and a translog function of labor and capital (as in Ackerberg et al., 2015). Given these assumptions, we describe our estimation procedure in detail in Online Appendix F. A possible concern with this estimation approach is that, if it were to be the case that materials are not perfect complements with the output of labor and capital, our measure of the firm's productivity would automatically incorporate a measure of the firm's materials' usage. This would be problematic for our identification approach, as firms may adjust their materials' usage directly in reaction to a demand-driven change in domestic sales. To address this possible concern, the second approach assumes instead that the production function is a Cobb-Douglas aggregator of materials and the same translog function of labor and capital employed in our baseline approach (see Bilir and Morales, 2018, for details on the estimation procedure). Thus, while our baseline approach imposes that material inputs have a zero elasticity of substitution with the output of labor and capital, the second approach imposes instead a unit elasticity of substitution.

In both estimation approaches, 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. (2016). Both estimation approaches do however use different outcome measures; while the approach that assumes a Leontief production function exploits data on the firm's sales revenue, the approach that assumes a Cobb-Douglas production function uses information on the firm's value added. We thus refer in Table 9 to the two measures of productivity that we obtain as "TFP Sales" and "TFP Value Added", respectively.

A general concern with our productivity estimates is that, if they do not correctly account 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

Table 9: Alternative TFP Measures

Dependent Variable:	ΔLn(Exports)			
	(1) OLS	(2) IV	(3) OLS	(4) IV
ΔLn(Domestic Sales)	-0.204[a] (0.027)	-1.602[a] (0.437)	0.105[a] (0.026)	-1.285[a] (0.486)
ΔLn(Avg. Wages)	-0.525[a] (0.051)	-1.149[a] (0.206)	-0.514[a] (0.064)	-0.873[a] (0.152)
ΔLn(TFP Sales)	1.357[a] (0.051)	2.657[a] (0.409)		
ΔLn(TFP Value Added)			0.807[a] (0.060)	1.218[a] (0.161)
F-Statistic		28.32		24.99
Observations	8,018	8,018	8,018	8,018

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality. All specifications include firm-level log average wages and sector and province fixed effects as additional controls.

point out that our measures of productivity are far from being perfectly correlated with the firm's total sales; specifically, this correlation is 0.55 for our baseline approach and 0.23 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.

A second concern with our productivity estimates is that, as we do not observe separately prices and quantities for each firm, they may capture no 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 evidence of this happening: the correlation between the boom-to-bust change in the number of vehicles per capita in the municipality of location of the firm and our productivity measures is very close to zero and its sign is actually negative (it is -0.04 for our baseline approach and -0.11 for our alternative approach).

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 slightly lower in absolute value than in our baseline specification (see column 2), we cannot reject the null hypothesis that this elasticity is equal to the baseline estimate of -1.602.

6.6   Placebo Tests of First-Stage Results

One could be concerned that the estimated impact of our instrument on firms' domestic sales (see columns 1 to 4 in Table 2) may due to the presence of underlying trends in economic conditions that affect both of these two variables and that are heterogeneous across municipalities in Spain. To evaluate the plausibility of this concern, we present in Table 10 the results of two related

Table 10: Placebo Tests of First Stage

Dependent Variable:	ΔLn(Domestic Sales)			
Sample:	Boom firms		Bust firms	
	Within Boom	Boom vs. Bust	Within Bust	Boom vs. Bust
	(1)	(2)	(3)	(4)
ΔLn(Vehicles p.c. in Municipality)	-0.041	0.184[b]	-0.009	0.277[a]
	(0.080)	(0.074)	(0.071)	(0.070)
Observations	5,344	5,344	5,245	5,245
F-statistic	0.27	6.19	0.02	15.63

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include firm-level log TFP and log average wages as additional controls. These coefficients are not included to save space. All specifications also include sector and province fixed effects. The sample use to compute the estimates in columns 1 and 2 includes all firms active in at least one year in the subperiod 2002-05 and in at least one year in the subperiod 2006-08. The sample use to compute the estimates in columns 3 and 4 includes all firms active in at least one year in the subperiod 2009-11 and in at least one year in the subperiod 2012-13.

falsification tests. In column 1, we break the boom period into two subperiods, 2002-05 and 2006-08, and evaluate whether our instrument (changes in demand between the boom and the bust periods) predicts the changes in domestic sales across these two subperiods. In column 3, we perform a similar exercise but with a dependent variable that measures changes in domestic sales between two subperiods included in the bust period, 2009-11 and 2012-13. If the correlation between changes in domestic sales and changes in the stock of vehicles per capita documented in the first four columns of Table 2 was due to underlying trends, the lack of synchronization between the time frames at which the endogenous variable and the instrument are measured should not affect the capacity of the latter to predict the former. However, as the results in columns 1 and 3 of Table 10 illustrate, the effect of our instrument in these placebo exercises is not statistically different from zero (the statistic of the F-test is below 0.3 in both cases) and, furthermore, the point estimates we obtain are very close to zero.

A possible concern with the interpretation of these placebo results is that they might be driven by the fact that the sample of firms in each of these placebo tests is different than in our baseline regressions (see notes to Table 10 for details). With that in mind, in columns 2 and 4 of Table 10 we repeat our baseline specifications comparing the boom (2002-08) to the bust (2009-13) for both the change in domestic sales and in the instrument, but for the same sample of firms used to compute the estimates reported in columns 1 and 3, respectively. In both cases, we continue to find a positive and statistically significant effect of the instrument on domestic sales, but the effect is admittedly much stronger in column 4 (with an F-stat of 15.63) than in column 2 (with an F-stat of 6.19).

6.7 Additional Robustness Tests

We finally perform a number of additional robustness tests. To save space, we include the exact estimates in Appendix D and focus here on summarizing the main findings. First, we show in Table D.1 that our results are not affected when excluding multinational subsidiaries operating in Spain from our sample, when weighting observations by the number of years a firm is active in export

markets, or when defining the bust period as 2010-2013 or 2011-13 instead of 2009-13. Second, we experiment in Table D.2 with additional variants of our instruments in Tables 2 and 7. More specifically, holding the municipality-level population constant at its 2002 level when computing the number of vehicles per capita in each municipality in both the boom and the bust periods has a negligible effect on our estimates. We also show that vehicles per capita both at the municipality and at the province-level remain significant when including them simultaneously in the first stage regression, and that our construction-sector instrument delivers similar results when it is based on employment shares or turnover shares, rather than on wage bill shares.

7   Structural Interpretation and Quantification

There is an obvious tension between our empirical results suggesting a negative impact of demand-driven changes in domestic sales on changes in exports, and the theoretical framework we describe in Section 2 to organize our empirical analysis. In this section, we show how a simple extension of that framework incorporating non-constant marginal costs delivers insights consistent with our empirical results. We close this section by using the extended framework to provide a quantitative assessment of the importance of the domestic slump for the observed export miracle in Spain during the period 2009-13.

7.1   Structural Interpretation

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

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

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). In Appendix A.2, we develop a micro-foundation for the cost function in equation (14) in a model in which a firm's short-run production capacity is limited by a fixed factor and the elasticity of output with respect to the variable factors is below one. Under this micro-foundation, the parameter λ is decreasing in the elasticity of output with respect to the variable factor, and $\lambda = 0$ when this elasticity is equal to one.

Solving for the optimal level of exports by firm i, and taking log differences, our model delivers

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

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

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

Solving for $\ln(\tau_d Q_{id} + \tau_x Q_{ix})$, plugging this expression into equation (15), and imposing the same decomposition as in equation (5), we then find that:

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

where $\varepsilon_{ix} \equiv u_{ix}^\xi + ((\sigma - 1)/(1 + \lambda))(u_i^\varphi - u_i^\omega)$. Note that this equation is analogous to the estimating equation (10) suggested in Section 2, except that it features the log difference of *total* sales (and not just domestic sales) on the right-hand side. 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 inclusion of the log change in exports as part of one of the right-hand-side variables generates an additional mechanical upward bias when estimating $(\sigma - 1)\lambda/(1 + \lambda)$. Note, however, that our instrumental variable approach continues to deliver consistent estimates of this coefficient provided that the only way that the instrument affects exporting is by affecting a firm's change in domestic sales and not by affecting exporting directly. In other words, as long as the instrument satisfies the exclusion restriction, estimating equation (17) via our instrumental variable approach should deliver consistent estimates of $(\sigma - 1)\lambda/(1 + \lambda)$.

In Table 11, we present OLS and two-stage least squares estimates of equation (17). In columns 1 to 3, we use the same productivity proxy φ_i^* as in previous tables. In columns 4 and 5, we additionally control for the initial stock of capital, as indicated by the micro-foundation in Appendix A.2. As expected, the OLS estimates in column 1 indicate a strong positive correlation between exports and total sales, even when including all the controls and fixed effects in equation (17). The first-stage results in column 2 indicate that our baseline instrument is a strong predictor of a firm's total sales (not just its domestic sales), with an F-stat of 23.42. Finally, the second-stage elasticity of exports to total sales in column 3 is negative and significant and stands at a value of -2.624. Adding the firm-specific stock of capital in the bust period does not affect significantly the first-stage nor the second-stage results. Thus, henceforth, we treat the estimates in column 3 of Table 11 as our baseline estimates.

To understand the magnitude of our estimates of the elasticity of exports with respect to domestic-demand-driven changes in total sales, take a firm with an initial export share of 16.5% (which corresponds to median export share during the boom in our sample of 8,018 continuing exporters). Suppose that, due to a drop in demand, this firm experiences a 10% drop in its domestic sales. Our estimated elasticity of exports to domestic sales in Table 2 indicates that, other things equal, the firm should see its exports increase by 16%. This also implies that the firm's total sales will decrease by $83.5\% \times 10\% + 16.5\% \times (-16\%) = 5.71\%$. Our estimated elasticity of exports to total sales in Table 11 then suggests an implied increase in exports of $5.71\% \times 2.624 = 14.98\%$, which is quite close to 16%. This demonstrates that our IV results in Tables 2 and 11 deliver congruent estimates for the response of exports to local demand shocks.[49]

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[49]In Online Appendix H, we present results for specifications analogous to those in Tables 1 to 10, with the only difference that the boom to bust log change in total sales is included as right-hand-side variable instead of the corresponding log change in domestic sales. The conclusions discussed in Section 6 are generally corroborated by the results shown in Online Appendix H.

Table 11: Intensive Margin with Total Sales

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
Δ Ln(Total Sales)	0.785^a (0.037)		-2.624^a (0.916)		-2.771^a (0.962)
Δ Ln(TFP)	0.397^a (0.053)	0.985^a (0.024)	3.741^a (0.900)	0.931^a (0.024)	3.670^a (0.894)
Δ Ln(Avg. Wages)	-0.095^b (0.047)	-0.432^a (0.031)	-1.567^a (0.408)	-0.416^a (0.030)	-1.563^a (0.413)
Δ Ln(Vehicles p.c. in Municipality)		0.221^a (0.046)		0.213^a (0.044)	
Δ Ln(Stock of Capital) (in bust period)				0.120^a (0.008)	0.481^a (0.121)
F-statistic		23.42		23.18	
Observations	8,018	8,018	8,018	8,018	8,018

Note: Standard errors clustered by municipality. All regressions include sector and province fixed effects. Significance levels: ^{a}p<0.01, ^{b}p<0.05, *p<0.1.

With an estimate of the demand elasticity σ in hand, it is easy to infer an estimated value of λ from the estimates in Table 11. Specifically, given the estimates in column 3, we can compute an estimate of λ as $\hat{\lambda} = 2.624/(\sigma - 2.624)$. For $\sigma = 6$ and $\sigma = 5$, we obtain $\hat{\lambda} = 0.77$, and $\hat{\lambda} = 1.10$, respectively, which in both cases indicates a significant departure from constant marginal costs.

7.2 Quantification

In this final section, we attempt 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 do so, we implement a three-step procedure. First, we measure for each sector s the boom-to-bust changes in the Spain-wide aggregate domestic demand shifter $Q_{sd} \equiv E_{sd}/P_{sd}$. Second, we compute the impact of several counterfactual boom-to-bust changes in Q_{sd} on the aggregate domestic, foreign, and total sales of Spanish firms; we define the counterfactual changes in Q_{sd} of interest as fractions of the observed changes in this aggregate shifter. Third, we perform a variance decomposition of the observed change in firms' total sales with the aim of informing the degree to which the boom-to-bust observed changes in Q_{sd} are truly due to demand or supply shocks. In the remainder of this section, we describe each of these three steps in detail. Finally, we conclude the section describing the results of our quantification.

Measuring sector-specific changes in aggregate demand shifter. Defining an aggregate demand shifter for sector s as $Q_{sd} \equiv E_{sd}/P_{sd}$, we can rewrite the demand that any firm i faces in Spain in any given period as $Q_{id} = (P_{id}/P_{sd})^{-\sigma} Q_{sd} \xi_{id}^{\sigma-1}$, where P_{sd} is the sectoral price index,

$$ P_{sd} = \left[\int_{i \in \mathcal{D}_s} \left(\frac{P_{id}}{\xi_{id}} \right)^{1-\sigma} di + \int_{i \in \mathcal{X}_s} \left(\frac{P_{id}}{\xi_{id}} \right)^{1-\sigma} di \right]^{\frac{1}{1-\sigma}}, $$

and $\mathcal{D}_s$ and $\mathcal{X}_s$ denote, respectively, the set of sector s domestic and foreign firms selling in Spain.

To measure the boom-to-bust change in Q_{sd} for every sector s, we first impose an equivalence

between the total expenditure in Spain in sector s, E_{sd}, and the sum of both the total domestic sales of all firms located in Spain and classified in sector s, $R_{sd} \equiv \int_{i \in \mathcal{D}_s} R_{id}$, and the Spanish aggregate expenditure on imported goods in sector s, $R_{sd}^{\mathcal{X}}$; i.e. $E_{sd} = R_{sd} + R_{sd}^{\mathcal{X}}$. Thus, we can write the boom-to-bust change in the sector s aggregate demand shifter as

$$\frac{Q_{sd1}}{Q_{sd0}} = \left(\frac{R_{sd1} + R_{sd1}^{\mathcal{X}}}{R_{sd0} + R_{sd0}^{\mathcal{X}}}\right) \Big/ \left(\frac{P_{sd1}}{P_{sd0}}\right),$$

where, for any variable x, we denote as x_1 and x_0 their respective boom and bust values. We measure R_{sd1} and R_{sd0} by aggregating the domestic sales of all firms in our dataset, $R_{sd1}^{\mathcal{X}}$ and $R_{sd0}^{\mathcal{X}}$ directly from the aggregate statistics on imports published by the Spanish Custom Agency, and P_{sd1}/P_{sd0} as the change in the Spanish Consumer Price Index.

Computing counterfactual changes in aggregate domestic sales and exports. Given our measure Q_{sd1}/Q_{sd0}, we define each of the counterfactual changes in Q_{sd} whose impact on firms' exports and domestic sales we study as

$$\frac{Q'_{sd1}}{Q_{sd0}} = \Gamma \frac{Q_{sd1}}{Q_{sd0}} + (1 - \Gamma), \tag{18}$$

where, for every variable x, we use x'_1 to denote its counterfactual value in the bust period, and $1 - \Gamma$ denotes the assumed contribution of demand shocks to the observed change in the aggregate sectoral demand shifter. We can thus interpret the counterfactual in which we set the change in the aggregate demand shifter to equal the expression in equation (18) as a counterfactual exercise that predicts how firms' aggregate domestic sales, exports, and total sales, would have changed if, for every sector, we had eliminated the demand-driven component of the observed change in Q_{sd}.[50]

When computing the aggregate change domestic sales and exports that we would have observed if the change in the demand shifter had been equal to Q'_{sd1}/Q_{sd0}, we 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. Specifically, we use data on the observed changes in domestic sales and exports of every Spanish firm to proxy for the boom-to-bust actual changes in the functions of these supply and idiosyncratic demand parameters that are relevant for our counterfactual exercise.

When implementing our counterfactual exercise, we assume that Spain is a small open economy and, thus, impose that counterfactual changes in the Spanish aggregate demand shifter do not affect: (a) the boom-to-bust change in the foreign price index P_{sx} and aggregate demand shifter Q_{sx}; (b) foreign firms' marginal production costs. Assumption (a) implies that Spain is a small exporter to the rest of the world; assumption (b) implies that Spain is a small importer from the rest of the world.

Given our results regarding the extensive margin of trade in Table 3, we assume that firms do not change their export status in reaction to our counterfactual change in the domestic demand shifter and, thus, focus on computing the effect of such counterfactual change on firms that either exported in both the boom and the bust periods, or that did not export in either period.

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[50]Although, according to our model, changes in Q_{sd} determine changes in the residual demand function that each firm faces and, thus, from the perspective of each individual firm, are purely demand shifters, these changes in Q_{sd} may be due, in general equilibrium, to either changes in productivity of the average manufacturing firm (i.e. supply shocks) or to changes in the propensity to consume of the average consumer (i.e. demand shocks). By setting it equal to $1 - \Gamma$, equation (18) implicitly imposes the assumption that the relative contribution of demand shocks to the change in Q_{sd} is the same for all sectors.

Given a value of Q'_{sd1}/Q_{sd0}, available on exports and domestic sales in boom and bust periods for every firm, and values of the demand parameter σ, and of the composite parameter that determines the within-firm elasticity of market-specific sales with respect to an exogenous change in total sales, $-((\sigma-1)\lambda)/(1+\lambda)$, the variant of the model described in Section 7.1 allows us to compute the counterfactual change in domestic sales R'_{id1}/R_{id0} for every firm in the economy, and the counterfactual change in exports R'_{ix1}/R_{ix0} for every firm that exports a positive amount in the boom and in the bust periods (see Appendix A.3 for details). We then aggregate these firm-specific counterfactual changes to constructed counterfactual changes in aggregate domestic sales and exports, R'_{d1}/R_{d0} and R'_{x1}/R_{x0}.

In our baseline specification, we set σ to 5, which is a central value in the range of estimates used in the international trade literature (see Head and Mayer, 2014), and we set $((\sigma-1)\lambda)/(1+\lambda)$ to the estimated value of 2.624 (see Table 11).[51]

To perform our counterfactual analysis, we derive three sets of equations from the model described in Section 7.1. The first set computes the counterfactual change in exports of every firm i belonging to a sector s and that exports in both boom and bust periods:

$$\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] = \ln\left[\frac{R_{ix1}}{R_{ix0}}\right] - \frac{(\sigma-1)\lambda}{1+\lambda}\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \qquad (19)$$

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. The second set of equations computes the counterfactual change in domestic sales of every firm i that belongs to a sector s and that is active in both boom and bust periods:

$$\ln\left[\frac{R'_{id1}}{R_{id0}}\right] = \ln\left[\frac{Q'_{sd1}}{Q_{sd0}}\left(\frac{Q_{sd1}}{Q_{sd0}}\right)^{-1}\right] + \sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right] + \ln\left[\frac{R_{id1}}{R_{id0}}\right]$$
$$- \frac{(\sigma-1)\lambda}{1+\lambda}\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \qquad (20)$$

where

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

denote the counterfactual change (relative to the actual change) in the aggregate sectoral demand shifter and price index, respectively. Finally, the system of counterfactual equilibrium equations includes an equation that yields the counterfactual change (relative to the actual change) in the sectoral price index:

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[51] As discussed in Online Appendix F, our production function estimation approaches also allows us to recover sector-specific estimates of σ. Depending on the period and the estimation procedure we use, we obtain average of these estimates across sectors that oscillate between 3 and 4. We opt to perform our counterfactual exercise under the calibrated value $\sigma = 5$ because our estimates crucially depend on a firm-specific measure of total variable costs. In our context, this measure is likely subject to substantial measurement error as our dataset lacks information on firms': (a) total energy expenditure; (b) separate average wage measures for temporary and permanent workers.

$$\ln\left[\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right] = \ln\left(s_{sd0}^{\mathcal{D}}\sum_{i\in\mathcal{D}_s} s_{id0}^{\mathcal{D}}\frac{R'_{id1}}{R_{id0}} + (1-s_{sd0}^{\mathcal{D}})\frac{Q'_{sd1}}{Q_{sd0}}\left(\frac{Q_{sd1}}{Q_{sd0}}\right)^{-1}\left(\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right)^{\sigma}\frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}}\right)$$

$$-\ln\left(s_{sd0}^{\mathcal{D}}\sum_{i\in\mathcal{D}_s}s_{id0}^{\mathcal{D}}\frac{R_{id1}}{R_{id0}} + (1-s_{sd0}^{\mathcal{D}})\frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}}\right) - \ln\left[\frac{Q'_{sd1}}{Q_{sd0}}\left(\frac{Q_{sd1}}{Q_{sd0}}\right)^{-1}\right], \tag{21}$$

where $s_{sd0}^{\mathcal{D}} \equiv R_{sd0}/(R_{sd0} + R_{sd0}^{\mathcal{X}})$ denotes the share of total expenditure in sector s that goes to Spanish firms, $s_{id0}^{\mathcal{D}} \equiv R_{id0}/R_{sd0}$ denotes firm i's share of total domestic sales of Spanish firms, and $R_{sd1}^{\mathcal{X}}/R_{sd0}^{\mathcal{X}}$ denotes the observed boom-to-bust change in Spanish imports in sector s.

The counterfactual boom-to-bust changes in total exports and domestic sales predicted by our model appear in Figure 6. For the set of firms that we use in our counterfactual analysis, aggregate domestic sales dropped 15.92% between the boom and the bust periods and exports grew by 11.99%. These are the values that our counterfactual analysis correctly generates when we set $1 - \Gamma = 0$: if the domestic slump had been entirely supply driven, then the vent-for-surplus mechanism would not have been operative and boom-to-bust changes aggregate exports and domestic sales would not be affected by a counterfactual that eliminates the contribution of demand shocks to the observed change the sectoral aggregate demand shifters, Q_{sd1}/Q_{sd0}. Conversely, if the change in the aggregate demand shifter had been entirely due to demand shocks, $1 - \Gamma = 1$, our model predicts that, in the absence of such demand shocks, aggregate domestic sales and aggregate exports would have dropped by 0.48% and 1.20%, respectively.

Figure 6: Impact of "Vent-for-Surplus" on Aggregate Domestic Sales and Exports
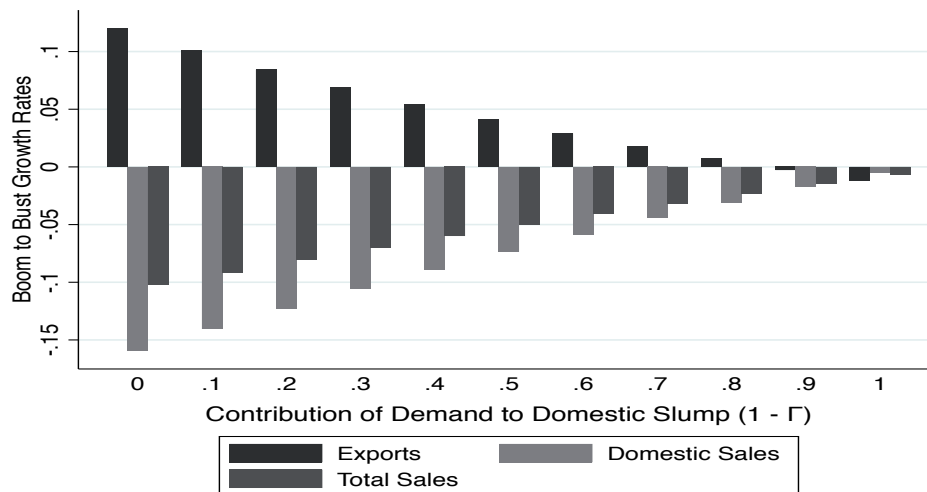
Notes: The horizontal axis indicates the value of the parameter $1 - \Gamma$. Given a value of $1 - \Gamma$, the export and domestic sales growth rates indicated in the vertical axis correspond to those predicted by equations (18) to (21).

Determining the contribution of the "vent-for-surplus" mechanism. Figure 6 indicates the predicted counterfactual growth rates in aggregate exports, domestic sales, and total sales for several different values of the parameter $1 - \Gamma$, which determines the relative contribution of demand shocks to the boom-to-bust observed change in the sectoral aggregate demand shifters Q_{sd1}/Q_{sd0}. To determine the empirically relevant value of this parameter, we perform a decomposition of the cross-firm variance of the observed boom-to-bust changes in total sales. Specifically, on the basis of equation (17), we decompose the variance of $\Delta \ln(R_i)$, with $R_i \equiv R_{id} + R_{ix}$, into a variance compo-

nent due to firms' marginal cost and export demand shifters and a variance component attributed to factors orthogonal to these shifters (see Appendix A.4 for details). When performing this decomposition, 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%.[52]

The procedure we follow to estimate the contribution of demand to the domestic slump is not without limitations. The variance decomposition that we implement reveals that 41% of the variance of the changes in firms' total sales is due to *any* residual factor that is orthogonal to firms' marginal cost shifters and export demand shocks. Thus, we can conclude that 41% of the variance in the firm-specific changes in total sales is due to demand factors if and only if we assume that these demand factors are orthogonal to the firms' marginal cost shifters and export demand shocks. If this assumption were not to hold, our 41% measure would capture the relative contribution to the variance of the changes in firms' total sales of only those demand components that happen to be orthogonal to the firms' marginal cost and export demand shifters. In this case, our 41% measure would be a lower bound on the contribution of demand shocks to the variance of the boom-to-bust firm-specific changes in total sales.

Quantification results. Depending on whether the observed changes in the aggregate demand shifter, Q_{sd1}/Q_{sd0} were entirely supply driven (i.e. $1 - \Gamma = 0$) or entirely demand driven (i.e. $1 - \Gamma = 1$), we would have observed a 9.60 percentage points difference (i.e. 10.23%-0.63%) in the boom-to-bust drop in aggregate total sales of Spanish firms. From our variance decomposition, one may infer that, in the absence of demand shocks, the boom-to-bust drop in aggregate total sales would have been $10.23\% + 41\% \times (0.63\% - 10.23\%) = 6.29\%$. This drop in aggregate total sales is very similar to that predicted by our model when $1 - \Gamma = 0.4$. We thus infer that the empirically relevant value of the contribution of demand to the domestic slump, $1 - \Gamma^*$, is close to 0.4. At this value of the parameter $1 - \Gamma$, our model predicts that, in the absence of demand shocks, the growth in exports would have been 5.43%. Given that the observed growth in exports was 11.99% (for $1 - \Gamma = 0$), our analysis indicates that the vent-for-surplus mechanism explains $(11.99\%\text{-}5.43\%)/11.99\% = 54.7\%$ of the total drop in exports.[53]

Looking at other outcomes of our counterfactual analysis, our model also predicts that, in the absence of any change in demand shocks between boom and bust periods, the total drop in domestic sales would have been equal to 8.91%; equivalently, our model indicates that these demand shocks explain $(15.91\%\text{-}8.91\%)/15.91\% = 44\%$ of the drop in domestic sales.

Finally, our model illustrates that exporters and non-exporters would have been affected differently by the change in the domestic demand shifters; in the absence of these changes, the exporters' aggregate domestic sales would have dropped by *only* 5.96% (in comparison to an observed dropped of 12.91%), while that of non-exporters would have dropped by 19.04% (in comparison to an observed dropped of 26.21%). In relative terms, our model predicts thus that demand shifters explain 53.83% and 27.16% of the observed dropped in the aggregate domestic sales of exporters and non-exporters, respectively. This difference between exporters and non-exporters in their drop in

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[52]If we were to first residualize $\Delta \ln(R_i)$ from the set of fixed effects and observed covariates included in equation (17), $var(\varepsilon_{ix})$ would explain 65% of the variance of the residualized values of $\Delta \ln(R_i)$. Factors orthogonal to ε_{ix} would thus explain 35% of the boom-to-bust variation in the residualized log changes in domestic sales.

[53]As we show in Appendix C.6, when using an estimate of $-((\sigma-1)\lambda)/(1+\lambda)$ equal to -1.819 (which is the largest estimate among those arising from the robustness exercises presented in Online Appendix H), our analysis indicates that the vent-for-surplus mechanism explains 44.1% of the total increase in exports.

domestic sales reflects that the former were much less affected by the negative marginal costs shocks (changes in productivity and wages) than the non-exporting firms.

8 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 benefitted from their freed capacity and consequent reduction in marginal production costs to increase their sales in foreign markets. We circumvent 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 mark-ups and technologies with constant marginal costs of production. We however 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.

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. With access to data on firms' exports and prices by destination market, one may potentially expand our analysis to a multi-country environment featuring a rich set of market-specific extensive margin and mark-up decisions. The interdependencies studied in this paper will naturally carry over to that environment, complicating the estimation of some key parameters of multi-country export models. However, we are hopeful that the tools in De Loecker et al. (2016), Antràs et al. (2017), Arkolakis and Eckert (2017), and Morales et al. (2018) will help surmount these complications and allow researchers to further study the role of interdependencies in shaping the response of firms to shocks to the world economy.

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A Theoretical Appendix

A.1 Biases Due to Measurement Error

We discuss here the implications of measurement error in both total sales and exports whenever domestic sales are computed by subtracting exports from the total sales of firms (see also, Berman et al., 2015).

Suppose that one does not observe R_{id} directly, but instead infers it from $R_i - R_{ix}$, where R_i is used to denote the total sales of a firm. Assume furthermore that both $\Delta \ln R_i$ and $\Delta \ln R_{ix}$ are measured with error, so that

$$
\begin{aligned}
\Delta \ln R_i &= \Delta \ln \breve{R}_i + \varpi_i \\
\Delta \ln R_{ix} &= \Delta \ln \breve{R}_{ix} + \varpi_{ix},
\end{aligned}
$$

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 can reach an estimating equation analogous to equation (10)

$$
\Delta \ln R_{ix} = d_s + d_\ell + (\sigma - 1)\, \delta_\varphi \Delta \ln(\varphi_i^*) - (\sigma - 1)\, \delta_\omega \Delta \ln(\omega_i^*) + \beta \Delta \ln R_{id} + \varepsilon_{ix},
$$

but we now have

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

Similarly, the error term in the expression for the change in domestic sales is given by

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

It then follows that the probability limit of the OLS estimator of the coefficient on domestic sales can we written as

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

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

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

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

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 controlling for (or partialling out) sector and location 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_i^{φ}, factor costs, u_i^{ω}, and export demand u_{ix}^{ξ}; and (c) it is mean independent of the measurement error in exports ϖ_{ix}.

A.2 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 seeking to produce a total amount of output Q_i can be expressed as:

$$\min \quad \omega_i L_i$$
$$s.t. \quad \varphi K_i^{\alpha_K} L_i^{\alpha_L} \geq Q_i,$$

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

$$\omega_i = \mu \alpha_L \frac{Q_i}{L_i}$$
$$\varphi K_i^{\alpha_K} L_i^{\alpha_L} = Q_i,$$

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

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

Using $\widetilde{\varphi}_i$ to denote a shifter of short-run marginal costs and λ to measure deviations of the short-run marginal cost function from a linear benchmark,

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

we can rewrite short-run 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 marginal costs function 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 marginal cost function. The curvature of the marginal cost schedule is thus crucially shaped by the reciprocal of the output elasticity of the flexible factor.

A.3 System of Equations for Counterfactual Exercise

In order to perform our counterfactual exercises, the two key equations from our model are:

$$\ln\left[\frac{R_{ix1}}{R_{ix0}}\right] = \ln\left[\frac{Q_{sx1}}{Q_{sx0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right]$$

$$+ \sigma\ln\left[\frac{P_{sx1}}{P_{sx0}}\right] - \frac{(\sigma-1)\lambda}{1+\lambda}\ln\left[\frac{R_{i1}}{R_{i0}}\right] \tag{22a}$$

$$\ln\left[\frac{R_{id1}}{R_{id0}}\right] = \ln\left[\frac{Q_{sd1}}{Q_{sd0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right]$$

$$+ \sigma\ln\left[\frac{P_{sd1}}{P_{sd0}}\right] - \frac{(\sigma-1)\lambda}{1+\lambda}\ln\left[\frac{R_{i1}}{R_{i0}}\right] \tag{22b}$$

where $\ln[x_1/x_0] \equiv \Delta\ln x$ denotes the log change between the boom and the bust periods in any covariate x, and remember that $R_{it} = R_{ixt} + R_{idt}$ for both $t=0$ and $t=1$. Equation (22a) is implied by equations (15) and (16), and equation (22b) is analogous for the case of the domestic market.

For any variable x, we define as x_1' the counterfactual value that this variable takes in the bust period if the value that the aggregate domestic demand shifter takes in the bust period Q_{sd1}' and all other demand and supply shocks had changed between boom and bust periods as they actually did. Therefore, analogously to equations (22a) and (22b), we can define the following two equations

$$\ln\left[\frac{R_{ix1}'}{R_{ix0}}\right] = \ln\left[\frac{Q_{sx1}}{Q_{sx0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right]$$

$$+ \sigma\ln\left[\frac{P_{sx1}}{P_{sx0}}\right] - \frac{(\sigma-1)\lambda}{1+\lambda}\ln\left[\frac{R_{i1}'}{R_{i0}}\right] \tag{23a}$$

$$\ln\left[\frac{R_{id1}'}{R_{id0}}\right] = \ln\left[\frac{Q_{sd1}'}{Q_{sd0}}\right] + (\sigma-1)\ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right] + \frac{(\sigma-1)}{1+\lambda}\left[\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right] - \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right]\right] - (\sigma-1)\ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right]$$

$$+ \sigma\ln\left[\frac{P_{sd1}'}{P_{sd0}}\right] - \frac{(\sigma-1)\lambda}{1+\lambda}\ln\left[\frac{R_{i1}'}{R_{i0}}\right] \tag{23b}$$

Note that equations (23a) and (23b) allow us to compute the impact of counterfactual demand shocks $\ln[Q_{sd1}'/Q_{sd0}]$ 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 (23b) applies for all these firms.

The assumption that the foreign price index is not affected by aggregate demand shocks in Spain is consistent with Spain being a small country relative to the foreign one. Domestic equilibrium wages would generally be affected by counterfactual changes in the aggregate demand term E_{sd}. Holding them at their observed path is however consistent with our aim of identifying the effect of demand shocks on exports working exclusively through changes in the aggregate demand shifter Q_{sd}. In other words, by omitting from our analysis the impact that, in general equilibrium, demand shocks may have had on equilibrium wages, we focus on the impact that these demand shocks have on firms' marginal production costs whenever these firms move along their marginal cost curve, but without accounting for any shift in this curve.

From equations (23a) and (23b), it is easy to see that the counterfactual changes in firm i's exports, $\ln[R_{ix1}'/R_{ix0}]$, and domestic sales, $\ln[R_{id1}'/R_{id0}]$, is a function of the actual changes in its own supply shocks and idiosyncratic demand shocks

$$\left\{\ln\left[\frac{\varphi_{i1}}{\varphi_{i0}}\right], \ln\left[\frac{\omega_{i1}}{\omega_{i0}}\right], \ln\left[\frac{\tau_{sd1}}{\tau_{sd0}}\right], \ln\left[\frac{\tau_{sx1}}{\tau_{sx0}}\right], \ln\left[\frac{\xi_{ix1}}{\xi_{ix0}}\right], \ln\left[\frac{\xi_{id1}}{\xi_{id0}}\right]\right\} \tag{24}$$

and, through the counterfactual change in the domestic price index, $\ln[P'_{sd1}/P_{sd0}]$, of the actual changes in the supply and idiosyncratic demand changes of all other firms that sell in the domestic market.

Using the expression for equilibrium exports and domestic sales predicted by the model described in Section 7.1, we can rewrite equations (23a) and (23b) in such a way that observed changes in exports and domestic sales are used to measure the impact of the different elements listed in equation (24) on the counterfactual change in exports and domestic sales. Specifically, we can rewrite equation (23a) as

$$\ln\left[\frac{R'_{ix1}}{R_{ix0}}\right] = \ln\left[\frac{R_{ix1}}{R_{ix0}}\right] - \frac{(\sigma-1)\lambda}{1+\lambda}\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \qquad (25)$$

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. Similarly, we can rewrite equation (23b) as

$$\ln\left[\frac{R'_{id1}}{R_{id0}}\right] = \ln\left[\frac{Q'_{sd1}}{Q_{sd0}}\left(\frac{Q_{sd1}}{Q_{sd0}}\right)^{-1}\right] + \sigma\ln\left[\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right] + \ln\left[\frac{R_{id1}}{R_{id0}}\right]$$
$$- \frac{(\sigma-1)\lambda}{1+\lambda}\left[\ln\left(\frac{R'_{ix1}}{R_{ix0}}\chi_{i0} + \frac{R'_{id1}}{R_{id0}}(1-\chi_{i0})\right) - \ln\left[\frac{R_{i1}}{R_{i0}}\right]\right], \qquad (26)$$

where

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

denote the counterfactual change (relative to the actual change) in the aggregate sectoral demand shifter and price index, respectively.

Besides the set of counterfactual changes in exports, R'_{ix1}/R_{ix0}, and domestic sales, R'_{id1}/R_{id0}, of every firm i located in Spain, the additional unknown in the system of equations formed by equations (25) and (26) is the counterfactual change in the domestic price index, P'_{sd1}/P_{sd0}. In order to understand how the domestic price index reacts to the supply and idiosyncratic demand changes of all the firms that sell in the domestic market, it is useful to rewrite this price index in any period t as

$$P_{sdt} = \frac{E_{sdt}}{Q_{sdt}} = \frac{R_{sdt} + R^{\mathcal{X}}_{sdt}}{Q_{sdt}}, \qquad (27)$$

where R_{sdt} denotes the total domestic sales of firms located in country d and operating in sector s, and $R^{\mathcal{X}}_{sdt}$ denotes the total 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

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

or, equivalently,

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

Simplifying notation, we can write that

$$\frac{P_{sd1}}{P_{sd0}} = \left(s_{sd0}^{\mathcal{D}} \frac{R_{sd1}}{R_{sd0}} + (1 - s_{sd0}^{\mathcal{D}}) \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}} \right) \left(\frac{Q_{sd1}}{Q_{sd0}} \right)^{-1},$$

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

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

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

$$\ln \left[\frac{P_{sd1}'}{P_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}} \right)^{-1} \right] = \ln \left(s_{sd0}^{\mathcal{D}} \sum_{i \in \mathcal{D}_s} s_{id0}^{\mathcal{D}} \frac{R_{id1}'}{R_{id0}} + (1 - s_{sd0}^{\mathcal{D}}) \frac{(R_{sd1}^{\mathcal{X}})'}{R_{sd0}^{\mathcal{X}}} \right)$$

$$- \ln \left(s_{sd0}^{\mathcal{D}} \sum_{i \in \mathcal{D}_s} s_{id0}^{\mathcal{D}} \frac{R_{id1}}{R_{id0}} + (1 - s_{sd0}^{\mathcal{D}}) \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}} \right) - \ln \left[\frac{Q_{sd1}'}{Q_{sd0}} \left(\frac{Q_{sd1}}{Q_{sd0}} \right)^{-1} \right], \qquad (28)$$

A key element in this expression is the variable $(R_{sd1}^{\mathcal{X}})'/R_{sd0}^{\mathcal{X}}$, 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

$$\frac{(R_{sd1}^{\mathcal{X}})'}{R_{sd0}^{\mathcal{X}}} = \frac{(R_{sd1}^{\mathcal{X}})'}{R_{sd0}^{\mathcal{X}}} \left[\frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}} \right]^{-1} \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}} = \frac{\frac{\sum_{i \in \mathcal{X}_s} R_{id1}'}{\sum_{i \in \mathcal{X}_s} R_{id0}}}{\frac{\sum_{i \in \mathcal{X}_s} R_{id1}}{\sum_{i \in \mathcal{X}_s} R_{id0}}} \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}} = \frac{\sum_{i \in \mathcal{X}_s} \left(\frac{R_{id0}}{\sum_{i \in \mathcal{X}_s} R_{id0}} \right) \frac{R_{id1}'}{R_{id0}}}{\sum_{i \in \mathcal{X}_s} \left(\frac{R_{id0}}{\sum_{i \in \mathcal{X}_s} R_{id0}} \right) \frac{R_{id1}}{R_{id0}}} \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}}$$

$$= \frac{\sum_{i \in \mathcal{X}_s} s_{id0}^{\mathcal{X}} \frac{R_{id1}'}{R_{id0}}}{\sum_{i \in \mathcal{X}_s} s_{id0}^{\mathcal{X}} \frac{R_{id1}}{R_{id0}}} \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}} = \frac{\sum_{i \in \mathcal{X}_s} s_{id0}^{\mathcal{X}} \frac{P_{id1}' Q_{id1}'}{P_{id0} Q_{id0}}}{\sum_{i \in \mathcal{X}_s} s_{id0}^{\mathcal{X}} \frac{P_{id1} Q_{id1}}{P_{id0} Q_{id0}}} \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}},$$

where $s_{id0}^{\mathcal{X}}$ 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

$$\frac{(R_{sd1}^{\mathcal{X}})'}{R_{sd0}^{\mathcal{X}}} = \frac{\sum_{i \in \mathcal{X}_s} s_{id0}^{\mathcal{X}} \frac{P_{id1}}{P_{id0}} \frac{Q_{id1}'}{Q_{id0}}}{\sum_{i \in \mathcal{X}_s} s_{id0}^{\mathcal{X}} \frac{P_{id1}}{P_{id0}} \frac{Q_{id1}}{Q_{id0}}} \frac{R_{sd1}^{\mathcal{X}}}{R_{sd0}^{\mathcal{X}}}, \qquad (29)$$

and we can write

$$\frac{Q'_{id1}}{Q_{id0}} = \left(\frac{P_{id1}}{P_{id0}}\right)^{-\sigma} \frac{Q'_{sd1}}{Q_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}, \tag{30}$$

$$\frac{Q_{id1}}{Q_{id0}} = \left(\frac{P_{id1}}{P_{id0}}\right)^{-\sigma} \frac{Q_{sd1}}{Q_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}, \tag{31}$$

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 (30) and (31) into equation (29), we can further rewrite the expression for the counterfactual change in Spanish imports in sector s as

$$\frac{(R^{\mathcal{X}}_{sd1})'}{R^{\mathcal{X}}_{sd0}} = \frac{\sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \frac{Q'_{sd1}}{Q_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}}{\sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \frac{Q_{sd1}}{Q_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}},$$

$$= \frac{\frac{Q'_{sd1}}{Q_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma} \sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}}{\frac{Q_{sd1}}{Q_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma} \sum_{i\in\mathcal{X}_s} s^{\mathcal{X}}_{id0} \left(\frac{P_{id1}}{P_{id0}}\right)^{1-\sigma} \left(\frac{\xi_{id1}}{\xi_{id0}}\right)^{\sigma-1}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}},$$

$$= \frac{\frac{Q'_{sd1}}{Q_{sd0}} \left(\frac{P'_{sd1}}{P_{sd0}}\right)^{\sigma}}{\frac{Q_{sd1}}{Q_{sd0}} \left(\frac{P_{sd1}}{P_{sd0}}\right)^{\sigma}} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}.$$

Plugging this expression into equation (28), we obtain an implicit equation for the change in the price index:

$$\ln\left[\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right] = \ln\left(s^{\mathcal{D}}_{sd0} \sum_{i\in\mathcal{D}_s} s^{\mathcal{D}}_{id0} \frac{R'_{id1}}{R_{id0}} + (1-s^{\mathcal{D}}_{sd0}) \frac{Q'_{sd1}}{Q_{sd0}} \left(\frac{Q_{sd1}}{Q_{sd0}}\right)^{-1} \left(\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1}\right)^{\sigma} \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}\right)$$

$$- \ln\left(s^{\mathcal{D}}_{sd0} \sum_{i\in\mathcal{D}_s} s^{\mathcal{D}}_{id0} \frac{R_{id1}}{R_{id0}} + (1-s^{\mathcal{D}}_{sd0}) \frac{R^{\mathcal{X}}_{sd1}}{R^{\mathcal{X}}_{sd0}}\right) - \ln\left[\frac{Q'_{sd1}}{Q_{sd0}}\left(\frac{Q_{sd1}}{Q_{sd0}}\right)^{-1}\right]. \tag{32}$$

Summing up, the relevant system of equations is given by equations (25), (26), and (32), which correspond to equations (19), (20), and (21) in the main text. The unknowns of this system are the counterfactual to actual relative change in the sectoral aggregate domestic price index,

$$\frac{P'_{sd1}}{P_{sd0}}\left(\frac{P_{sd1}}{P_{sd0}}\right)^{-1};$$

for every exporting firm in sector s, its counterfactual change in exports

$$\frac{R'_{ix1}}{R_{ix0}};$$

and, for every active firm (no matter whether it exports or not) in sector s, its counterfactual change in domestic sales

$$\frac{R'_{id1}}{R_{id0}}.$$

Every other element in the system formed by the equations (19), (20), and (21) is either estimated (as it is the case of the parameters σ and $(\sigma-1)\lambda/(1+\lambda)$) or is directly observed in the data. Once we have computed the counterfactual changes in exports and domestic sales for every exporter and every domestic firm in the economy, respectively, we compute their counterfactual exports and domestic sales in the bust, and add the resulting numbers to compute the aggregate growth rates in exports and domestic sales shown in Figure 6.

A.4 Decomposition of the Variance of Boom-to-Bust Changes in Total Sales

We can rewrite equation (17) as

$$\Delta \ln \mathcal{R}_{ix} = \beta \Delta \ln \mathcal{R}_i + \varepsilon_{ix}, \tag{33}$$

with

$$\beta = -\frac{(\sigma-1)\lambda}{1+\lambda} \qquad \text{and} \qquad \varepsilon_{ix} \equiv u_{ix}^{\xi} + \frac{(\sigma-1)}{1+\lambda}(u_i^{\varphi} - u_i^{\omega}), \tag{34}$$

where, as a reminder, we denote by $\Delta \ln \mathcal{X}$ the residual of a regression of a variable $\Delta \ln X$ on a set of sector fixed effects $\{d\}_s$, location fixed effects $\{d\}_\ell$, and the observable covariates $\Delta \ln \varphi_i^*$, and $\Delta \ln \omega_i^*$. Using this notation, we can write the probability limit of the OLS and IV estimators of β as

$$\beta_{ols} = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}_i)}{var(\Delta \ln \mathcal{R}_i)}, \qquad\qquad \beta_{iv} = \frac{cov(\Delta \ln \mathcal{R}_{ix}, \Delta \ln \mathcal{R}_i^*)}{cov(\Delta \ln \mathcal{R}_i, \Delta \ln \mathcal{R}_i^*)}, \tag{35}$$

with $\Delta \ln \mathcal{R}_i = \Delta \ln \mathcal{R}_i^* + \Delta \ln \mathcal{R}_i^{\varepsilon}$, and 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}, and $\Delta \ln \mathcal{R}_i^{\varepsilon}$ is the part of $\Delta \ln \mathcal{R}_i$ correlated with ε_{ix}. In practice, given an estimate $\hat{\beta}_{iv}$, we recover an estimate of ε_{ix} for every exporter i as $\Delta \ln \mathcal{R}_{ix} - \hat{\beta}_{iv}\Delta \ln \mathcal{R}_i$ and we compute an estimate of $\Delta \ln \mathcal{R}_i^{\varepsilon}$ by running a regression of $\Delta \ln \mathcal{R}_i$ on $\Delta \ln \mathcal{R}_{ix} - \hat{\beta}_{iv}\Delta \ln \mathcal{R}_i$.

After simple algebraic manipulations, we can relate β_{ols} and β_{iv} as

$$\beta_{ols} = \beta_{iv}\frac{var(\Delta \ln \mathcal{R}_i^*)}{var(\Delta \ln \mathcal{R}_i)} + \beta_{\varepsilon}\left(1 - \frac{var(\Delta \ln \mathcal{R}_i^*)}{var(\Delta \ln \mathcal{R}_i)}\right), \tag{36}$$

and, thus, we can compute the share of the variance in total sales that is due to factors orthogonal to the unobserved supply shocks u_i^{φ} and u_i^{ω} and export demand shocks u_{ix}^{ξ} as

$$\frac{var(\Delta \ln \mathcal{R}_i^*)}{var(\Delta \ln \mathcal{R}_i)} = \frac{\beta_{ols} - \beta_{\varepsilon}}{\beta_{iv} - \beta_{\varepsilon}}. \tag{37}$$

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%.

We also perform a similar analysis to that described in equations (33) to (37) but without previously controlling for any fixed effect or any proxy for the firms' marginal cost shifters. In this case, our procedure will yield a decomposition of the cross-firm variance in the observed changes in total sales, $var(\Delta \ln R_i)$, into a component that is due to the impact on $\Delta \ln R_i$ of variables correlated with the regression residual,

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

and a component that is due to the impact on $\Delta \ln(R_i)$ of variables that are orthogonal to $\tilde{\varepsilon}_{ix}$. 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 is not without concerns, as it requires assuming that our instrument is valid unconditionally, and not just conditionally on sector and location fixed effects and our proxies for firms' factor prices and productivity.

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 and Spain come from 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) and the Eurosystem. We use the input-output tables produced by the Spanish National Statistical Office (Instituto Nacional de Estadística) 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 4).

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 Comercial Central*), which contains annual financial statements of around 85% of registered firms in the non-financial economy. We collate data from two separate sources to construct our own firm-level dataset: (i) the *Central de Balances* dataset from Banco de España and (ii) SABI, from Informa, a private company. Despite being based on the same original source, these two datasets are complementary: the first includes the largest number of firms and has the best coverage of small and medium enterprises, while the second has the most precise coverage of large firms. A detailed description of how we combine the two sources to construct our firm-level dataset can be found in Almunia, Lopez-Rodriguez and Moral-Benito (2018).

B.3   Foreign Transactions Dataset

As described in Section 3.3, the Bank of Spain requires 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 onwards, 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 entry correspond 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 year 2008.

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   Instruments

We construct several instrumental variables using information available at the either zip code, municipality or province level. The information on the stock of vehicles by both municipality and province is provided by the Spanish Registry of Motor Vehicles (Dirección General de Tráfico). The data on the number of foreign tourists and foreign overnight stays at the province level come from the Spanish National Statistical Office (Instituto Nacional de Estadística). The information on the population by both municipality and province is provided by the Spanish National Statistical Office (Instituto Nacional de Estadística). The data to construct the proxy for the housing supply elasticity in a given municipality come from the Spanish Cadastre (Dirección General del Catastro). In particular, we use the measure in Basco and Lopez-Rodriguez (2017), which is a municipality-specific ratio of available "buildable" urban land to urban land with already built structures. The ratio is calculated in a year sufficiently removed from the housing boom (1996) to avoid feedback effects of booming prices on the availability of "buildable" urban land during the Spanish housing bubble in the 2000s. The information on the residential house prices at municipality level used in Appendix C 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 price indexes for residential housing are calculated from 2004 to 2012 for municipalities with more than 1000 inhabitants and more than 30 transactions per year. These indexes are deflated using the Consumer Price Index provided by the Spanish National Statistical Office (Instituto Nacional de Estadística).

C   Appendix Figures

C.1   Share of Exports and GDP Within the European Union

Figure C.1 plots the share of exports to non-EU countries and GDP for Greece, Portugal, Spain and Germany (see Appendix B.1 for information on the sources of data).

Figure C.1: Share of Exports to non-EU Countries and GDP

Panel (a): Greece

Panel (b): Portugal

Panel (c): Spain

Panel (d): Germany
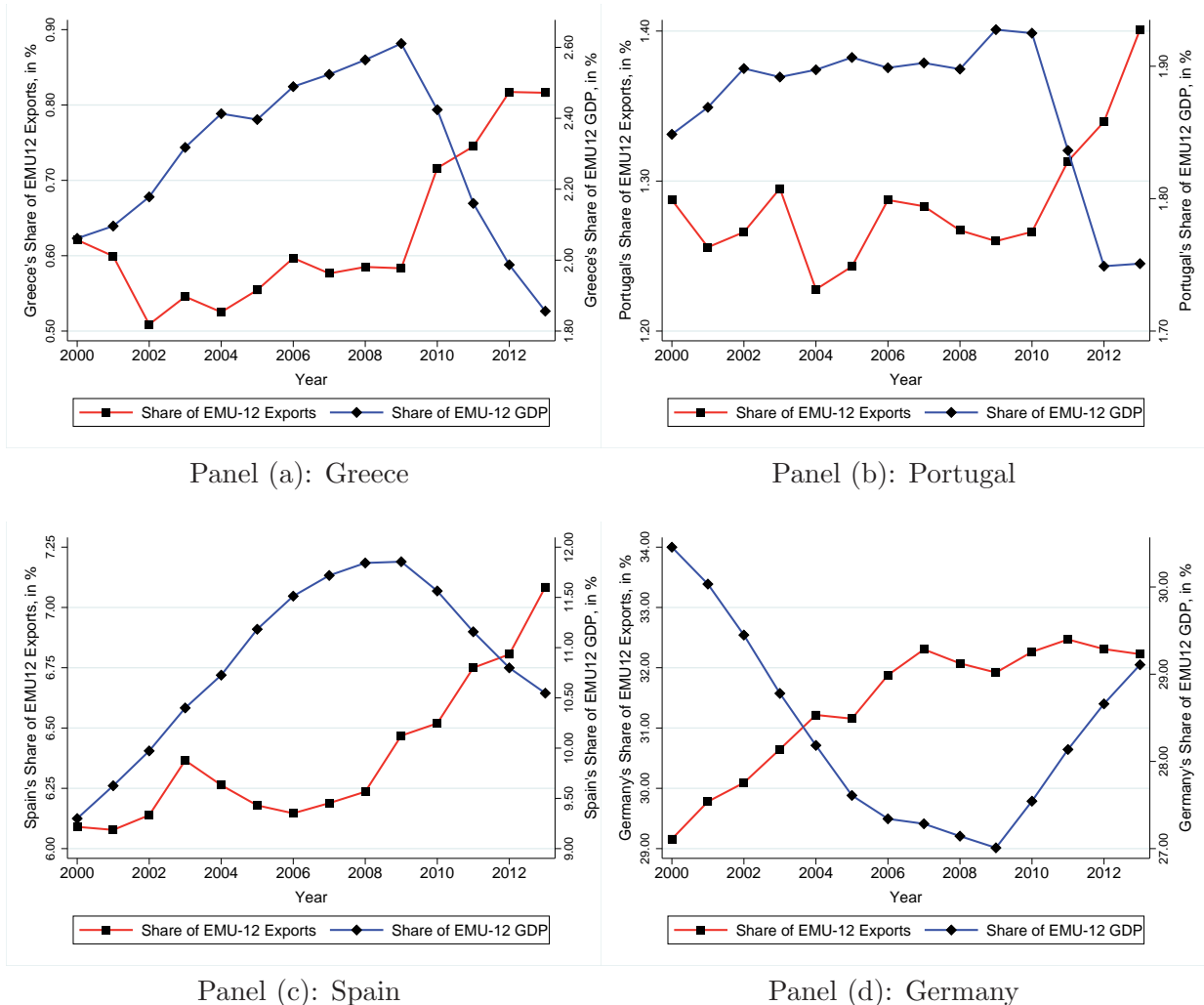
C.2   Spatial Distribution of Economic Activity in Spain

Figure C.2 plots the 2002-2008 annual average number of firms and number of exporting firms for each of the 47 Spanish peninsular provinces (see Appendix sections B.2 and B.3 for information on the sources of data).

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

(a) Number of Firms

(b) Number of Exporting Firms
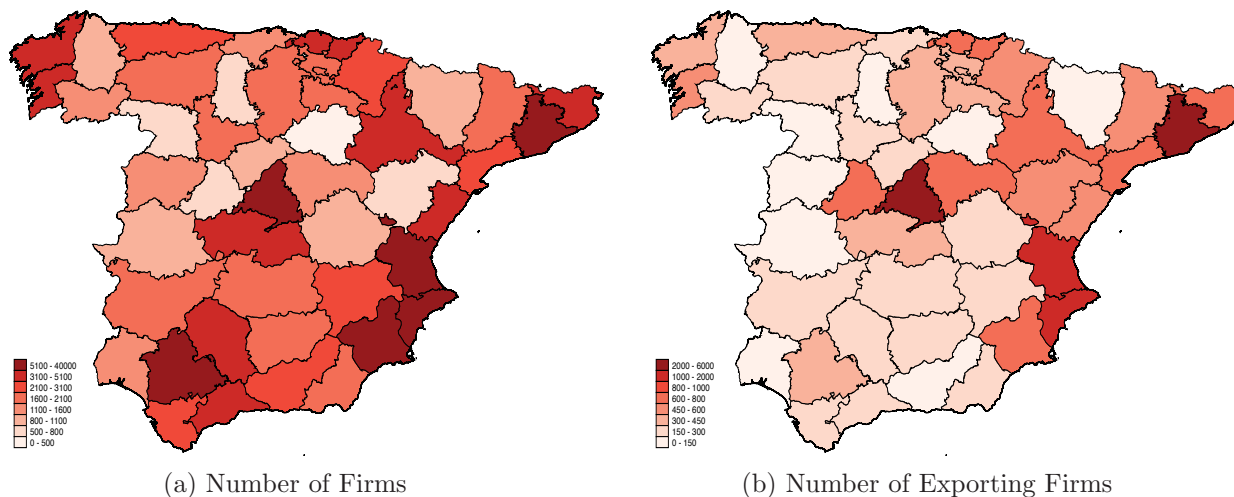
C.3 Province-Level Home Bias in Spanish Manufacturing

Data are for the year 2007 from C-interreg. We are grateful to Carlos Llano for providing them to us (see Llano et al., 2010, for details on this database).

Figure C.3: Province-Level Home Bias in Spanish Manufacturing
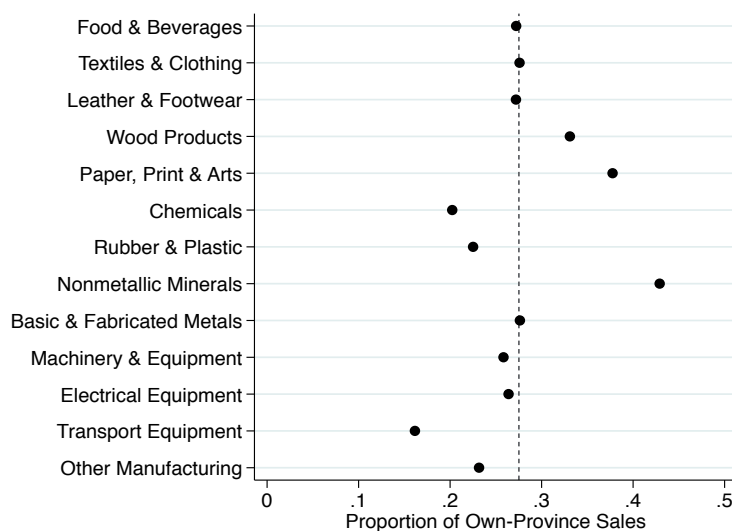
C.4 First-Stage and Reduced-Form Relationships

The two panels in Figure C.4 provide a graphical representation of the relationship between the boom-to-bust change in the log of the number of vehicles per capita in a municipality and the boom-to-bust change in the log of domestic sales (panel a) and exports (panel b) of the firms located in that municipality. Panel (a) thus represents the first-stage relationship between the endogenous covariate and the instrument, while panel (b) represents the reduced-form relationship between the outcome variable of interest and the instrument.

Figure C.4: First-Stage and Reduced-Form Relationships
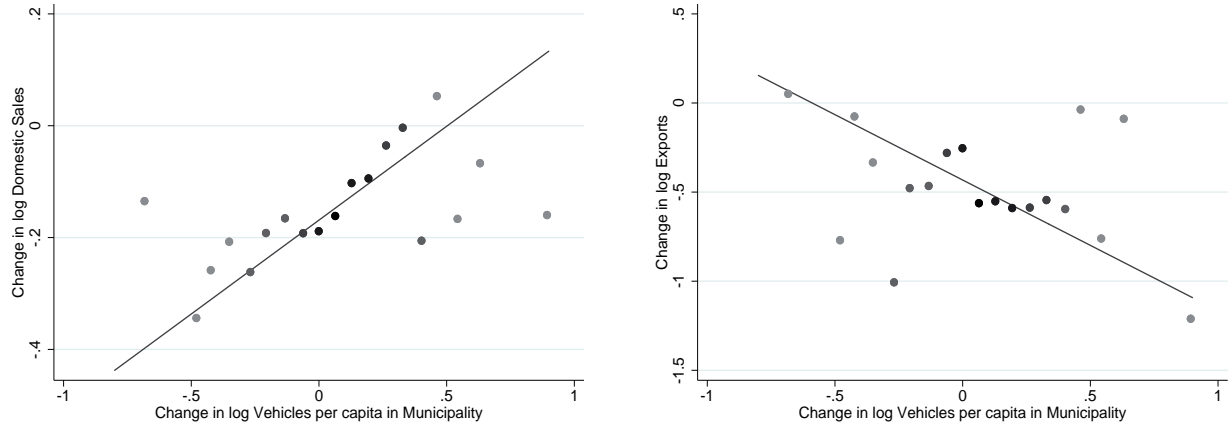
(a) First-Stage Relationship

(b) Reduced-Form Relationship

Notes: Each dot represents the average change in log real domestic sales (panel a) and in log real exports (panel b) for a given value of the change in log vehicles per capita in a municipality. 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=8,018) as in the regressions of Table 2, without including any controls or fixed effects.

C.5   Housing Supply Elasticities and Price Growth

The following figure shows that there is a negative correlation between the 1996 ratio of available 'buildable' urban land to urban land with already built structures (a proxy for the housing supply elasticity) and housing price growth between 2004 and 2007.

Figure C.5: Housing Supply Elasticities and Housing Price Growth during 2004-07
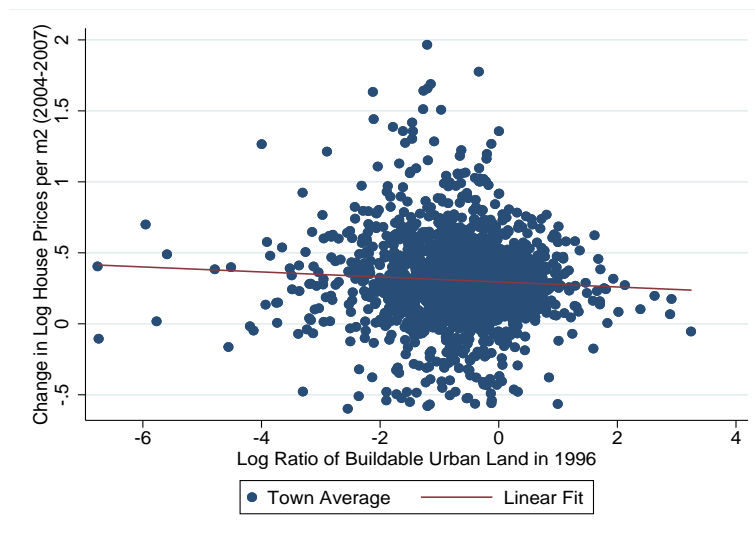
C.6 Counterfactual Export Growth Under Alternative Parametrizations

Figure C.6 is analogous to Figure 6. It differs from it only in that it is computed under an estimate of $-((\sigma-1)\lambda)/(1+\lambda)$ equal to -1.819 (instead of the baseline estimate of -2.624). This estimate of $-((\sigma-1)\lambda)/(1+\lambda)$ is the largest one among all the different estimates arising from the robustness exercises presented in Online Appendix H.

Given an estimate of $-((\sigma-1)\lambda)/(1+\lambda)$ equal to -1.819, we infer from our variance decomposition that, in the absence of demand shocks, the boom-to-bust drop in aggregate total sales would have been $10.23\% + 41\% \times (-0.01\% - 10.23\%) = 5.61\%$. Given this drop in domestic sales, we infer (as in our baseline quantification in Section 7.2) that the contribution of demand to the domestic slump, $1 - \Gamma^*$, was close to 0.4. At this value of the parameter $1 - \Gamma$, our model predicts that, in the absence of demand shocks, the growth in exports would have been 6.71%. Given that the observed growth in exports was 11.99%, our analysis indicates that the vent-for-surplus mechanism explains $(11.99\%\text{-}6.71\%)/11.99\% = 44.1\%$ of the total drop in exports.

Figure C.6: Impact of "Vent-for-Surplus" on Aggregate Domestic Sales and Exports
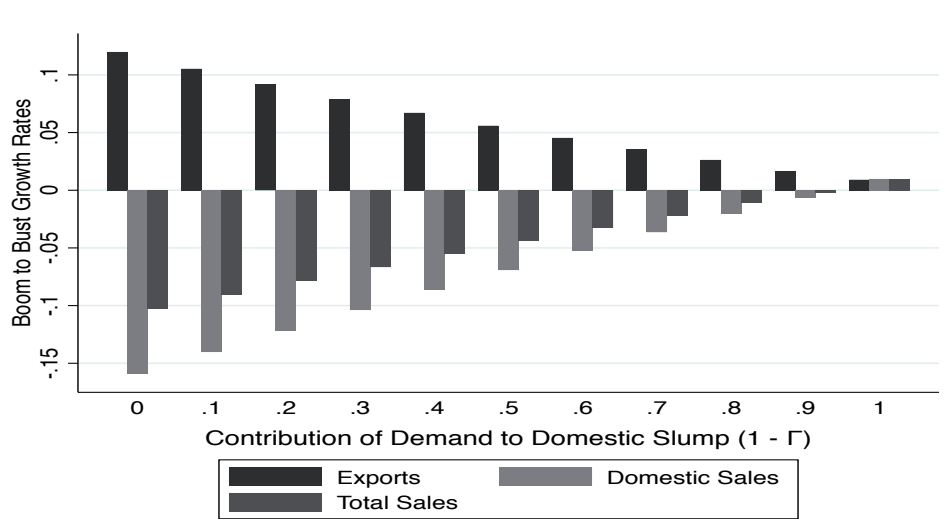
Notes: The horizontal axis indicates the value of the parameter $1 - \Gamma$. Given a value of $1 - \Gamma$, the export and domestic sales growth rates indicated in the vertical axis correspond to those predicted by equations (18) to (21).

D Appendix Tables

Table D.1: Additional Robustness Tests

Sample:	Excluding multinationals (1)	Weight by # of years exporting (2)	Bust as 2010-2013 (3)	Bust as 2011-2013 (4)
OLS Elasticity	-0.194^a	-0.185^a	-0.207^a	-0.241^a
	(0.030)	(0.025)	(0.029)	(0.031)
IV Elasticity	-1.518^a	-1.324^a	-1.648^a	-1.540^a
	(0.469)	(0.460)	(0.497)	(0.477)
1st Stage Coefficient	0.336^a	0.313^a	0.327^a	0.327^a
	(0.066)	(0.066)	(0.069)	(0.068)
Observations	6,629	8,018	7,357	6,722
F-statistic	26.11	22.54	22.32	22.89

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include sector and province fixed effects. Vehicles p.c. denotes the stock of vehicles per capita. First stage coefficient and F-statistic denote the corresponding statistic for the vehicles p.c. covariate.

Table D.2: Additional Alternative Instruments and Overidentification Tests

Dependent Variable:	ΔLn(Domestic Sales)			
	(1)	(2)	(3)	(4)
Δ Ln(Vehicles p.c. in in Municipality, 2002 pop.)	0.349^a (0.047)			
Δ Ln(Vehicles p.c. in Province)		0.561^b (0.271)		
Δ Ln(Vehicles p.c. in Municipality)		0.246^a (0.081)		
Δ Ln(Construction Employment) $\times$ 2002 Employment Share in Municipality			0.379^a (0.071)	
Δ Ln(Construction Turnover) $\times$ 2002 Turnover Share in Municipality				0.147^a (0.023)
F-statistic	54.60	19.70	28.31	41.54
	ΔLn(Exports)			
Δ Ln(Domestic Sales)	-1.677^a (0.336)	-1.285^a (0.326)	-1.349^b (0.587)	-1.929^a (0.423)
p-value for Sargan test	0.80	0.46	0.80	0.68
Observations	8,018	8,018	7,928	7,928

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality, except for column 2, where they are clustered by province. All specifications include firm-level log TFP and log wages as additional controls (coefficients not included to save space). All specifications include province and sector fixed effects, except column 2 which only includes sector effects.

Venting Out: Exports during a Domestic Slump

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

Online Appendix

E   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}. Omitting the subindex t from the notation for simplicity, we can thus express a dummy taking value one if firm i exports as

$$d_{ix} = \mathbb{1}\{\ln R_{ix} > \sigma \ln F_{ix}\}$$

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} that includes a set of period-specific sector fixed effects, location fixed effects, and observed proxies φ_i^* and ω_i^* is

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

Focusing on a linear probability model, we further rewrite the probability of the firm exporting as

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

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

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

where, from equation (6),

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

with the different terms in this expression defined as in Section 2, and ΔX_{ix} is a vector of sector fixed effects, location fixed effects and first-differences in the observed proxies φ_i^* and ω_i^*. We analogously decompose the log change in fixed costs of exporting as

$$\Delta \ln F_{ixt} = \phi_{sx} + \phi_{\ell x} + \phi_\varphi \Delta \ln(\varphi_i^*) + \phi_\omega \Delta \ln(\omega_i^*) + u_i^F,$$

similarly to how we decomposed the demand shifter, productivity and cost levels in Section 2. Notice that we are being quite flexible, letting firm-level fixed export costs depend on firm-level productivity and factor costs, and on both sector and location fixed effects. In particular, $\Delta \ln F_{ixt}$ depends on the same elements included in the vector ΔX_{ix} and the additional term u_i^F.

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

$$\Delta \Pr(\Delta X_{ix}) = \mathbb{E}\big[(\gamma_{sx} - \phi_{sx}) + (\gamma_{\ell x} - \phi_{\ell x}) + [(\sigma - 1)\gamma_\varphi - \sigma\phi_\varphi]\Delta \ln(\varphi_i^*)$$
$$- [(\sigma - 1)\delta_\omega - \sigma\phi_\omega]\Delta \ln(\omega_i^*) + \beta \Delta \ln R_{id} + \varepsilon_{ix} - \sigma u_i^F \big|\Delta X_{ix}\big],$$

where, as in equation (9), $\varepsilon_{ix} = (\sigma - 1)[u_{ix}^\xi + u_i^\varphi - u_i^\omega]$. Following the same steps as in Section 2, the following asymptotic properties of $\hat{\beta}_{OLS}$ can be derived:

$$plim(\hat{\beta}_{OLS}) = \frac{cov(u_{ix}^\xi + u_i^\varphi - u_i^\omega - \frac{\sigma}{\sigma - 1}u_i^F, u_{id}^\xi + u_i^\varphi - u_i^\omega)}{var(u_{id}^\xi + u_i^\varphi - u_i^\omega)}.$$

The only difference relative to equation (11) is the addition of the term $-u_i^F$ 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

$$plim(\hat{\beta}_{IV}) = \frac{cov(u_{ix}^\xi + u_i^\varphi - u_i^\omega - \frac{\sigma}{\sigma-1}u_i^F, \mathcal{Z}_{id})}{cov(u_{id}^\xi + u_i^\varphi - u_i^\omega, \mathcal{Z}_{id})}.$$

This expression will equal zero as long as the instrument $\widehat{\mathcal{Z}}_{id}$ verifies the following two conditions: (a) it is correlated with the boom-to-bust change in domestic sales of firm i, after controlling for (or partialling out) firm and location 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_i^φ, factor costs, u_i^ω, export demand shocks, u_{ix}^ξ, and export fixed-cost shocks u_i^F (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_{id}^ξ.

It is straightforward to extend the above analysis to the case in which total sales and exports are measured with error and domestic sales are imputed by subtracting exports from total sales. Following the same steps as in Appendix A.1, we obtain

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

and, thus, the only additional requirement on the instrument is that it is mean independent of the measurement error in exports ϖ_{ix}.

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 (2018). 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:

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

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

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

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 in this Appendix section, 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.2, 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, Fort and Tintelnot (2017). Given the assumption that all firms are monopolistically competitive in their output markets, it will be true that

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

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

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

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

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

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

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

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

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

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

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

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, 2018, for the corresponding estimation approach). No matter which of these two production functions we assume, we estimate the corresponding production function parameters and demand parameters separately for the boom and bust periods and for each of the twenty-four 2-digit NACE sectors in which the manufacturing firms in our dataset are classified. For both the boom and the bust periods, we report here the simple average across all sectors of the estimated labor and capital elasticities, of the estimated persistence parameters ρ and of the demand elasticity σ.

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

Under the assumption that the production function is Cobb-Douglas in materials, we obtain the following estimates. In the boom period, the average elasticities of value added with respect to

labor and capital are 0.76 and 0.19, respectively; the average annual autocorrelation in performance is 0.78; and the average demand elasticity is 3.19. In the bust period, the average elasticities of value added with respect to labor and capital are 0.86 and 0.17, respectively; the average annual autocorrelation in performance is 0.78; and the average demand elasticity is 3.05.

Notice that both estimation approaches yield estimates of the demand elasticity σ that are a bit low relative to those that, using identification strategies different from ours, are typically obtained in the international trade literature (see Head and Mayer, 2014, for a review). One possible explanation for this mismatch between our estimates and those in the international trade literature is the fact that we cannot observe firms' expenditure in energy; this may imply that our measure of the variable production costs underestimates the firms' total expenditure in variable inputs and, thus, that our estimates of σ are downward biased. These estimates of σ do not, however, impact 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 7.2. 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 (2018). 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.
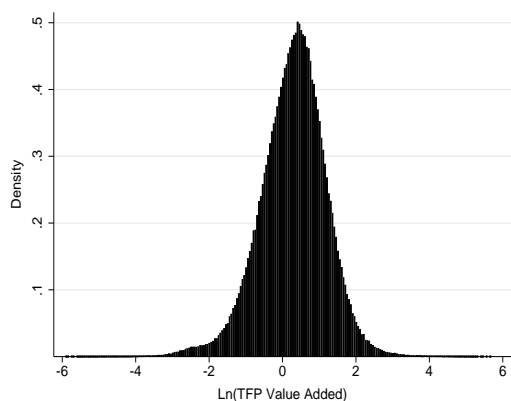
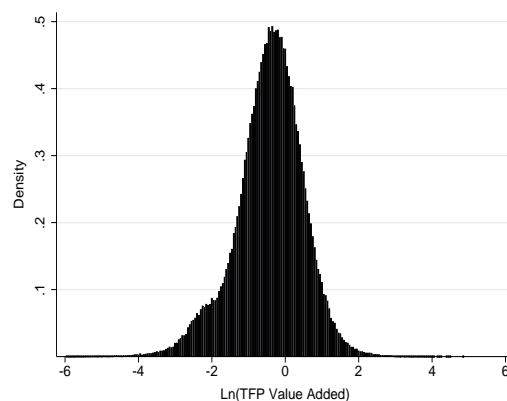## Figure F.1: Productivity Estimates
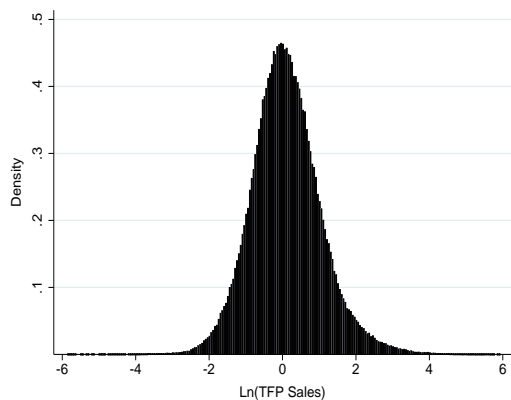
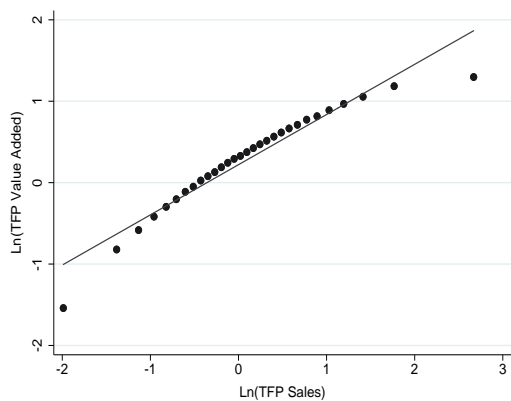(a) Density of Ln(TFP Sales) in Boom
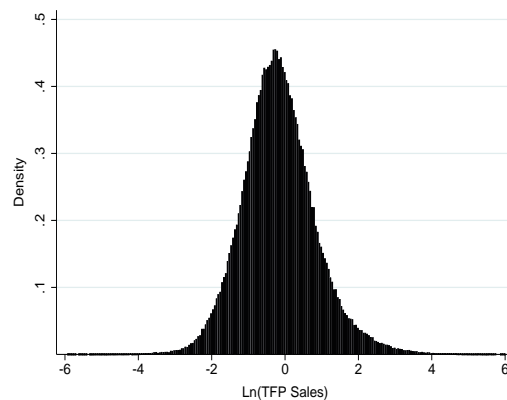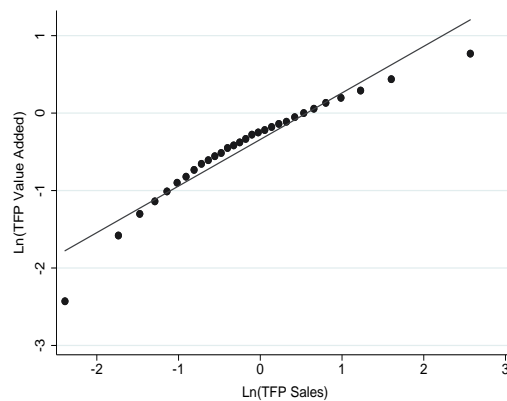

(b) Density of Ln(TFP Sales) in Bust

(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 (2018). 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.

G Exports and Domestic Sales: Year-to-Year Variation

In the main text we focus on empirical specifications that compare export behavior in a *bust* period relative to a *boom* period. In this Appendix section, we study OLS specifications that exploit year-to-year variation in domestic sales and exports. Before doing so, however, we extend the discussion in Section 2 to a setup with multiple periods.

G.1 Theoretical Model

We assume here that equations (1) to (3) hold in every year t and, thus, we can write the expression for log exports by firm i in year t as:

$$\ln R_{ixt} = \kappa + (\sigma - 1)\left[\ln \xi_{ixt} + \ln \varphi_{it} - \ln \omega_{it}\right] - (\sigma - 1)\left(\ln \tau_{xt} - \ln P_{xt}\right) + \ln E_{xt}, \qquad \text{(G.1)}$$

where κ is a constant. In order to transition into an estimating equation, we model the demand, productivity and cost levels as follows:

$$
\begin{aligned}
\ln(\xi_{ixt}) &= \xi_{ix} + \xi_{xt} + \tilde{\xi}_{ix} \times t + u^{\xi}_{ixt}, \\
\ln(\varphi_{it}) &= \varphi_i + \varphi_t + \tilde{\varphi}_i \times t + \delta_\varphi \ln(\varphi^*_{it}) + u^{\varphi}_{it}, \\
\ln(\omega_{it}) &= \omega_i + \omega_t + \tilde{\omega}_i \times t + \delta_\omega \ln(\omega^*_{it}) + u^{\omega}_{it}.
\end{aligned}
\qquad \text{(G.2)}
$$

Note that we are decomposing these terms into (i) a time-invariant firm fixed effect, (ii) a firm-invariant year fixed effect (which in the regressions will be expanded to include a whole set of municipality-year and sector-year fixed effects), (iii) a firm-specific linear trend, (iv) an observable part of these terms for the case of productivity (φ^*_{it}) and input bundle costs (ω^*_{it}), and (iv) a residual term on which we impose an assumption analogous to that in footnote 16. Given these decompositions, we can re-write equation (G.1) as:

$$\ln R_{ixt} = \kappa + \gamma_{ix} + \gamma_{xt} + \tilde{\gamma}_{ix} \times t + (\sigma - 1)\,\delta_\varphi \ln(\varphi^*_{it}) - (\sigma - 1)\,\delta_\omega \ln(\omega^*_{it}) + \varepsilon_{ixt}, \qquad \text{(G.3)}$$

where $\gamma_{ix} \equiv (\sigma - 1)\left[\xi_{ix} + \varphi_i - \omega_i\right]$, $\gamma_{xt} \equiv (\sigma - 1)\left[\xi_{xt} + \varphi_t - \omega_t\right] - (\sigma - 1)\left(\ln \tau_{xt} - \ln P_{xt}\right) + \ln E_{xt}$, $\tilde{\gamma}_{ix} \equiv (\sigma - 1)\left[\tilde{\xi}_{ix} + \tilde{\varphi}_i - \tilde{\omega}_i\right]$, and the error term is given by

$$\varepsilon_{ixt} = (\sigma - 1)\left[u^{\xi}_{ixt} + u^{\varphi}_{it} - u^{\omega}_{it}\right]. \qquad \text{(G.4)}$$

Following similar steps, we can derive the expression for revenues in the local market:

$$\ln R_{idt} = \kappa + \gamma_{id} + \gamma_{dt} + \tilde{\gamma}_{id} \times t + (\sigma - 1)\,\delta_\varphi \ln(\varphi^*_{it}) - (\sigma - 1)\,\delta_\omega \ln(\omega^*_{it}) + \varepsilon_{idt}, \qquad \text{(G.5)}$$

where $\gamma_{id} \equiv (\sigma - 1)\left[\xi_{id} + \varphi_i - \omega_i\right]$, $\gamma_{dt} \equiv (\sigma - 1)\left[\xi_{dt} + \varphi_t - \omega_t\right] - (\sigma - 1)\left(\ln \tau_{dt} - \ln P_{dt}\right) + \ln E_{dt}$, $\tilde{\gamma}_{id} \equiv (\sigma - 1)\left[\tilde{\xi}_{id} + \tilde{\varphi}_i - \tilde{\omega}_i\right]$, and

$$\varepsilon_{idt} = (\sigma - 1)\left[u^{\xi}_{idt} + u^{\varphi}_{it} - u^{\omega}_{it}\right]. \qquad \text{(G.6)}$$

Consider now using OLS to estimate the parameters of the following regression, which includes log domestic sales as an additional covariate in equation (G.3):

$$\ln R_{idt} = \kappa + \gamma_{id} + \gamma_{dt} + \tilde{\gamma}_{id} \times t + (\sigma - 1)\,\delta_\varphi \ln(\varphi^*_{it}) - (\sigma - 1)\,\delta_\omega \ln(\omega^*_{it}) + \beta \ln R_{idt} + \varepsilon_{idt}. \quad \text{(G.7)}$$

From equations (G.6) and (G.7), the probability limit of the ordinary least-squares (OLS) estimator of the coefficient on domestic sales can we written as

$$plim(\hat{\beta}_{OLS}) = \frac{cov(\ddot{r}_{ixt}, \ddot{r}_{idt})}{var(\ddot{r}_{idt})} = \frac{cov(u_{ixt}^\xi + u_{it}^\varphi - u_{it}^\omega, u_{idt}^\xi + u_{it}^\varphi - u_{it}^\omega)}{var(u_{idt}^\xi + u_{it}^\varphi - u_{it}^\omega)}, \qquad (G.8)$$

where we denote by $\ddot{X}$ the residual of a regression of a variable X on a set of firm fixed effects, year fixed effects, firm-specific linear time trends, and the proxies $\ln \varphi_{it}^*$ and $\ln \omega_{it}^*$. Analogously to the discussion in Section 2, it holds that:

1. As long as productivity and production factor costs are not perfectly observable or captured by the various fixed effects or the firm-specific trends, there will be a spurious positive correlation between exports and domestic sales that, in large samples, would lead one to estimate a positive value of $\hat{\beta}_{OLS}$ even when marginal costs are constant and, thus, demand-driven changes in domestic sales do not impact exports.

2. In the presence of a non-zero correlation in the residual demand (partialling out fixed effects) faced by firms in domestic and foreign markets, $\hat{\beta}_{OLS}$ will also converge to a non-zero value. Because this residual variation in demand does not capture market-specific macro shocks (which are controlled for via municipality-year and sector-year fixed effects), it seems plausible that u_{ixt}^ξ and u_{idt}^ξ will be positively correlated, leading one again to estimate a positive value of $\hat{\beta}_{OLS}$.

We have focused so far on the intensive margin, i.e., the impact of domestic demand shocks on the level of exports conditional on exporting. As in Appendix E, an analysis of the extensive margin of exports delivers very similar insights. More specifically, even if the true elasticity of the probability of exporting to demand-driven changes in domestic sales were to be 0, one is likely to estimate a positive elasticity whenever productivity and production factor costs are not perfectly captured by the various fixed effects, firm-specific trends and observable proxies, or whenever unobserved *residual* demand shocks are positively correlated across markets.

G.2 Empirical Results: Intensive Margin

Table G.1 presents OLS estimates from different specifications in which the logarithm of exports of a firm in a given year is regressed on the logarithm of its domestic sales in the corresponding year and different sets of controls. When no firm-specific controls are included in the regression, the discussion in Appendix section G.1 implies that we expect to observe a positive relationship between a firm's domestic sales in a given year and its volume of exports. This positive relationship is indeed observed in column 1 of Table G.1, in which we estimate an elasticity of export flows with respect to domestic sales of 0.645. The only controls in that regression are sector-year and municipality-year fixed effects.

In the remaining columns of Table G.1, we control for various sources of marginal cost and export demand heterogeneity across firms, with the aim of attenuating the biases identified in equation (G.8). In column 2, we introduce firm fixed effects, thus controlling for differences in firm characteristics that are constant over time and that may impact their productivity, factor prices and export demand shifters. The resulting estimated elasticity is very close to zero, -0.074, consistent with the predictions of the constant marginal cost model. Columns 3 and 4 additionally control for observed time-varying determinants of firms' marginal costs. Specifically, we control in column 3 for a measure of the firm's productivity (estimated, as in the main text, following the procedure in Gandhi et al., 2016), and we additionally control in column 4 for a measure of the firm's average wages (reported by the firm in its financial statement). Consistent with the results in the boom to bust regressions in the main text, controlling for these supply shocks reduces the OLS estimate of the coefficient on domestic sales, which goes down to -0.263. This indicates that, once we control for the firm's supply shocks, domestic sales and exports are negatively correlated. Columns 5 and

6 aim to additionally control for unobserved determinants of firms' marginal costs that are time varying. To do so, we additionally include firm-specific time trends as controls. The resulting estimates are again lower and indicate that a 10% decrease in a firm's domestic sales, keeping its productivity and average wages constant, implies a 3.19% increase in its aggregate export flows.

The last two columns in Table G.1 re-estimate the regression models in columns 2 and 6 using a specification in first-differences (instead of in levels). The differences between the coefficients on the domestic sales covariate in columns 2 and 7 (higher in the specification in levels than in that in first differences) reflect the fact that, while some of the missing covariates in these two specifications (i.e., firms' time-varying productivity and average wages) are strongly serially correlated and share common underlying trends with the corresponding firm's domestic sales, their year-to-year variation is less correlated with the yearly changes in domestic sales. Consistently with this interpretation, once we control for these serially correlated determinants of firms' marginal costs, the coefficient on domestic sales in the levels specification (column 6) becomes very similar to that in the first-differences specification (column 8). Given that the specifications in columns 6 and 8 yield very similar estimates, but the latter is computationally easier to estimate, we focus on the specification in first differences in the remaining tables presented in this Appendix.

As discussed in Section 5.1, one might be concerned that because total sales is a key input in the computation of our firm-level measure of TFP, our empirical results are just unveiling a mechanical negative correlation between exports and domestic sales once one holds total sale revenue constant. The correlation beween log TFP and log total sales in our data is however far from

Table G.1: Intensive Margin

Dependent Variable:			Ln(Exports)				ΔLn(Exports)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Ln(Domestic Sales)	0.645^a	-0.074^a	-0.231^a	-0.263^a	-0.228^a	-0.319^a		
	(0.013)	(0.017)	(0.016)	(0.015)	(0.016)	(0.016)		
Ln(TFP)			1.031^a	1.344^a		1.061^a		
			(0.042)	(0.048)		(0.063)		
Ln(Average Wages)				-0.627^a		-0.463^a		
				(0.041)		(0.047)		
ΔLn(Domestic Sales)							-0.228^a	-0.320^a
							(0.012)	(0.013)
ΔLn(TFP)								0.917^a
								(0.047)
ΔLn(Average Wages)								-0.408^a
								(0.033)
Observations	54,575	54,276	54,276	54,276	54,276	54,276	54,575	54,276
R-squared	0.474	0.898	0.904	0.906	0.951	0.953	0.237	0.368
Firm FE	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Sector-Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Municipality-Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm-specific trends	No	No	No	No	Yes	Yes	No	No

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by firm in parentheses. Exports, domestic sales and average wages are in constant 2011 euros. For any variable X, $\Delta Ln(X)$ is the difference in $Ln(X)$ between two consecutive years.

perfect. Specifically, the correlation between log changes in TFP and log total sales is 0.31. Notice also that columns 5 and 7 in Table G.1 unveil a negative correlation between log changes in exports and log changes in domestic sales even when *not* controlling for log changes in TFP.

In Figure G.2, we complement the analysis in Table G.1 by estimating sector-specific elasticities of exports with respect to domestic sales (see Table G.2 in Appendix D for the associated

regression tables).[1] The main conclusion is that the negative elasticity between domestic sales and exports documented in Table G.1 is pervasive across nearly all manufacturing sectors, the only exception being the "Pharmaceutical Products" sector, whose 95% confidence interval is such that we cannot reject the null hypothesis that, after controlling for firm-specific fixed effects and time trends and observed measures of productivity and labor costs, changes in domestic sales have no impact on exports. For all remaining sectors, the estimated elasticity of interest oscillates between -0.156 (manufacture of leather and related products) and -0.565 (manufacture of paper and paper products).

Table G.2: Intensive Margin - Heterogeneity by Sector

Dependent Variable:			ΔLn(Exports)				
Sector	Food	Beverages	Textiles	Clothing	Leather	Wood	Paper
NACE code	10	11	13	14	15	16	17
ΔLn(Domestic Sales)	-0.244^a	-0.325^a	-0.264^a	-0.447^a	-0.156^a	-0.234^a	-0.565^a
	(0.041)	(0.082)	(0.048)	(0.105)	(0.029)	(0.056)	(0.117)
ΔLn(TFP)	0.750^a	0.243	0.578^a	0.827^a	0.449^a	0.512^a	1.093^a
	(0.101)	(0.201)	(0.143)	(0.217)	(0.138)	(0.196)	(0.192)
ΔLn(Average Wages)	-0.350^a	0.091	-0.464^a	-0.179	-0.225^a	-0.003	-0.585^a
	(0.065)	(0.146)	(0.102)	(0.168)	(0.078)	(0.125)	(0.211)
Observations	7,578	1,729	2,489	968	2,620	1,725	1,434
R-squared	0.302	0.361	0.379	0.382	0.348	0.402	0.341

Sector	Printing	Chemicals	Pharma.	Plastic	Non-metals	Basic Metals	Fabric. Metals
NACE code	18	20	21	22	23	24	25
ΔLn(Domestic Sales)	-0.531^a	-0.357^a	-0.378^c	-0.456^a	-0.426^a	-0.241^a	-0.409^a
	(0.097)	(0.048)	(0.214)	(0.065)	(0.054)	(0.066)	(0.040)
ΔLn(TFP)	1.082^a	1.125^a	0.973^a	1.238^a	0.910^a	0.870^a	1.424^a
	(0.320)	(0.203)	(0.305)	(0.129)	(0.162)	(0.140)	(0.129)
ΔLn(Average Wages)	-0.620^a	-0.490^a	-0.448^c	-0.514^a	-0.271^b	-0.338^b	-0.493^a
	(0.224)	(0.112)	(0.250)	(0.104)	(0.118)	(0.133)	(0.090)
Observations	1,258	4,854	1,166	3,979	3,190	2,179	7,226
R-squared	0.343	0.286	0.305	0.345	0.388	0.325	0.283

Sector	Computers	Electron.	Machine	Vehicles	Furniture	Repair	Other
NACE code	26	27	28	29	31	32	33
ΔLn(Domestic Sales)	-0.321^a	-0.414^a	-0.280^a	-0.367^a	-0.452^a	-0.351^a	-0.369^a
	(0.082)	(0.058)	(0.024)	(0.048)	(0.092)	(0.075)	(0.062)
ΔLn(TFP)	1.023^a	0.911^a	1.114^a	0.992^a	1.227^a	1.262^a	1.562^a
	(0.249)	(0.259)	(0.146)	(0.202)	(0.231)	(0.309)	(0.399)
ΔLn(Average Wages)	-0.360^b	-0.430^c	-0.604^a	-0.545^a	-0.546^a	-0.557^a	-0.653^b
	(0.146)	(0.225)	(0.120)	(0.153)	(0.159)	(0.181)	(0.300)
Observations	1,490	2,307	7,502	2,833	1,407	1,287	1,211
R-squared	0.321	0.336	0.241	0.309	0.354	0.349	0.388

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. All specifications contain firm fixed effects and province-year fixed effects. Standard errors clustered by firm in parentheses. Exports, domestic sales and average wages are in constant 2011 euros. For any variable X, ΔLn(X) is the difference in Ln(X) between two consecutive years.

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[1] We exclude the "tobacco", the "petroleum refining", and the "other transport" sectors (two-digit industry codes 12, 19 and 30, respectively) due to the extremely low number of firms in those sectors.

Figure G.2: Intensive Margin - Heterogeneity by Sector
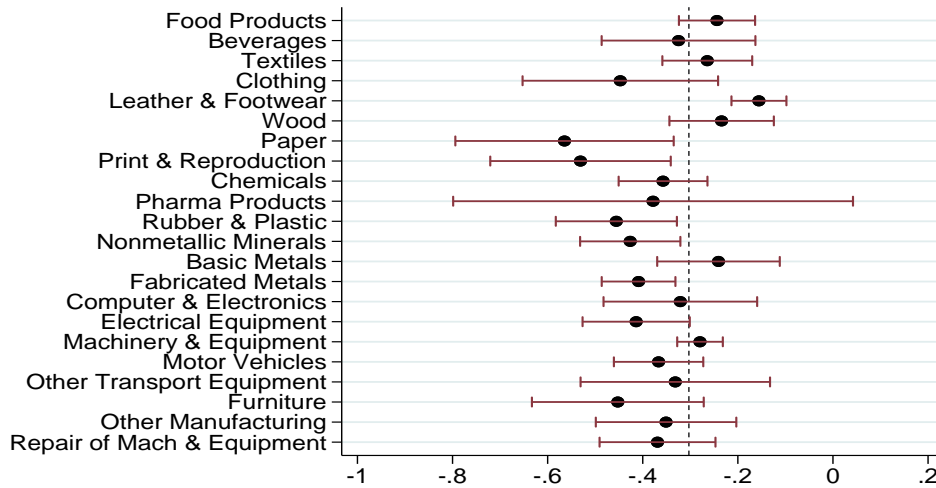
Note: The dotted vertical line reflects the average estimate reported in column 8 of Table G.1. The black dots reflect the sector-specific point estimates; the red lines reflect the 95% confidence interval for each of the sectoral estimates.

G.3 Extensive Margin

Local demand shocks may not only generate an intensive margin change in the export volume of those firms participating in export markets but may also lead firms to either start exporting or to stop participating in foreign markets, thus affecting the extensive margin of trade. We next explore the effect of demand-driven changes in domestic sales on the probability that a firm exports. To do so, we estimate three different types of binary choice models: static conditional logit models (columns 1 to 5 in Table G.3), a static linear probability models (column 6 in Table G.3), and a dynamic linear probability model (column 7 in Table G.3).

The results regarding the impact of domestic sales on the extensive margin of exports are similar to those described above for its impact on the intensive margin of exports. When we do not include any control for firm-specific marginal costs, we observe a positive correlation between a firm's domestic sales and its probability of exporting. As columns 2 and 3 in Table G.3 illustrate, controlling either for firm fixed effects only or for firm and sector-year fixed effects reduces the coefficient on domestic sales in absolute value but preserves its positive sign. Nevertheless, when we include controls for observable time-varying determinants of a firm's marginal cost, the elasticity of export participation with respect to domestic sales becomes negative. The most general conditional logit specification that we run accounts for firm fixed effects, sector-year fixed effects, and firm-year specific measures of productivity and average wages (column 5); the resulting elasticity of the export probability with respect to domestic sales is -0.312.[2]

Contrary to the specifications discussed in Appendix G.2, those discussed in columns 1 to 5 of Table G.3 do not account for firm-specific time trends. Accounting for both firm fixed effects and firm-specific time trends in a conditional logit model would be problematic for two reasons. First, it would be computationally very challenging. Second, it would give rise to an incidental parameters problem (Chamberlain, 1980), resulting in inconsistent estimates of the elasticity of

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[2]The parameters in columns 2 to 5 have been estimated following the procedure in Chamberlain (1980). This estimation procedure maximizes a conditional likelihood function that does not depend on the firm fixed effects and, consequently, does not yield estimates of these unobserved effects. However, given the nonlinear nature of the model, the elasticity of the export probability with respect to domestic sales does depend on these unobserved effects. For the exclusive purposes of computing the elasticities reported in the last row of Table G.3, we have set all these unobserved effects equal to zero.

Table G.3: Extensive Margin

Model:	Conditional Logit					Linear	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Ln(Domestic Sales)	1.184^a	0.141^a	0.255^a	-0.164^a	-0.312^a	-0.030^a	-0.038^a
	(0.008)	(0.018)	(0.020)	(0.024)	(0.026)	(0.001)	(0.001)
Log TFP				1.530^a	2.134^a	0.065^a	0.064^a
				(0.050)	(0.063)	(0.003)	(0.003)
Log Average Wages					-0.982^a	-0.033^a	-0.033^a
					(0.050)	(0.002)	(0.002)
Lagged Participation							0.161^a
							(0.006)
Observations	747,519	129,712	129,712	129,712	129,712	747,519	495,151
Firm FE	No	Yes	Yes	Yes	Yes	Yes	Yes
Sector-Year FE	Yes	No	Yes	Yes	Yes	Yes	Yes
Firm-specific trends	No	No	No	No	No	Yes	No
Elasticities	1.035	0.016	0.013	-0.150	-0.312	-0.234	-0.284

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Columns 1 to 5 present Maximum Likelihood estimators of the corresponding conditional logit model. Column 2 to 5 present estimates computed following the procedure in Chamberlain (1980). Column 6 presents OLS estimates of the corresponding linear probability model. Column 7 presents estimates computed following the procedure in Arellano and Bond (1991). The number of observations in columns 1, 6 and 7 correspond to the number of firm-years that we observe in our sample when taking into account all firms. The number of observations in columns 2 to 5 correspond to the number of firm-years that we observe in our sample when taking into account only those firms that change their export status at least once during the sample period. Standard errors in parentheses. Exports, domestic sales and average wages are in constant 2011 euros. For any variable X, $\Delta Ln(X)$ is the difference in $Ln(X)$ between two consecutive years.

export participation with respect to domestic sales.[3] Consequently, to test the robustness of our estimates to accounting for firm-specific time trends, we resort to the linear probability model specification. The estimates in column 6 of Table G.3 predict an elasticity of export participation with respect to domestic sales of -0.234, very similar to that predicted by the conditional logit model in column 5.

As shown in Das et al. (2007) and Morales at al. (2018), the export decision of firms is dynamic, depending both on their prior export status as well as on their expectations of future potential profits that a firm may earn by entering export markets. While correctly accounting for firms' expectations of future export profits is beyond the scope of this paper (see Dickstein and Morales, 2018), accounting for the prior export status of each firm only requires additionally controlling for a dummy that captures each firm's one-year lagged export participation (see Roberts and Tybout, 1997). We introduce this control in column 7 of Table G.3: the resulting estimate of the export participation elasticity with respect to domestic sales is -0.284, very similar to those obtained in columns 5 and 6.

In sum, our OLS estimates in Tables G.1, G.2, and G.3 demonstrate the existence of a strong, within-firm negative relationship between demand-driven changes in domestic sales and changes in the intensive and extensive margin of exports. As discussed in Section 2, however, it is reasonable to expect that these OLS estimates underestimate the extent to which demand-driven reductions in domestic demand generate expansions in export markets.

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[3]Charbonneau (2017) introduces a new estimator that allows to consistently estimate binary logit models in the presence of an individual-specific fixed effect and a choice-specific fixed effect. This estimator does not apply to our context, in which both sets of unobserved effects are firm-specific.

H   Regression Results with Total Sales instead of Domestic Sales

We present in this Appendix section specifications analogous to those in tables 1 to 10, with the only difference that the boom to bust log change in total sales is included as right-hand-side variable instead of the corresponding log change in exports.

Table H.1: Intensive Margin: Ordinary Least Squares Estimates

Dependent Variable:			ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)
Δ Ln(Total Sales)	0.891^a	0.778^a	0.764^a	0.778^a	0.785^a	0.801^a
	(0.029)	(0.036)	(0.037)	(0.039)	(0.037)	(0.041)
Δ Ln(TFP)		0.302^a	0.334^a	0.409^a	0.397^a	0.382^a
		(0.046)	(0.053)	(0.055)	(0.053)	(0.057)
Δ Ln(Avg. Wages)			-0.069	-0.107^b	-0.095^b	-0.051
			(0.043)	(0.047)	(0.047)	(0.050)
Observations	8,018	8,018	8,018	8,018	8,018	7,507
R-squared	0.164	0.170	0.170	0.210	0.223	0.327
Sector FE	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	No	No	No	No	Yes	No
Municipality FE	No	No	No	No	No	Yes

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, ΔLn(X) is the difference in Ln(X) between its average in the 2009-2013 period and its average in the 2002-2008 period. The estimation sample includes all firms selling in at least one year in the period 2002-2008 and in the period 2009-2013.

Table H.2: Intensive Margin: Two-Stage Least Squares Estimates

Dependent Variable:	ΔLn(Domestic Sales)				ΔLn(Exports)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ Ln(Total Sales)					-4.767^b	-2.220^a	-2.326^a	-2.624^a
					(2.355)	(0.770)	(0.763)	(0.916)
Δ Ln(Vehicles p.c. in Municipality)	0.154^a	0.271^a	0.262^a	0.221^a				
	(0.052)	(0.047)	(0.045)	(0.046)				
Δ Ln(TFP)		0.827^a	0.983^a	0.985^a		2.771^a	3.361^a	3.741^a
		(0.020)	(0.024)	(0.024)		(0.642)	(0.762)	(0.900)
Δ Ln(Avg. Wages)			-0.519^a	-0.432^a			-1.675^a	-1.567^a
			(0.026)	(0.031)			(0.419)	(0.408)
F-statistic	8.84	33.07	34.34	23.42				
Observations	8,018	8,018	8,018	8,018	8,018	8,018	8,018	8,018
Sector FE	No	No	No	Yes	No	No	No	Yes
Province 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 municipality 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. Vehicles p.c. denotes the stock of vehicles per capita. 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(Vehicles p.c. in Municipality) equals zero.

Table H.3: Extensive Margin: Two-Stage Least Squares Estimates

Dependent Variable:		Export Dummy		Proportion of Years	
	1st Stage (1)	OLS (2)	2nd Stage (3)	OLS (4)	2nd Stage (5)
Ln(Total Sales)		0.072^a	-0.133	0.054^a	-0.088
		(0.003)	(0.225)	(0.002)	(0.119)
Ln(Vehicles p.c. in Municipality)	0.072^a				
	(0.022)				
Ln(TFP)	1.091^a	0.002	0.226	0.014^a	0.168
	(0.015)	(0.005)	(0.247)	(0.003)	(0.130)
Ln(Average Wages)	-0.415^a	-0.010^b	-0.096	-0.017^a	-0.076
	(0.011)	(0.004)	(0.094)	(0.002)	(0.050)
F-statistic	10.20				
Observations	125,808	125,808	125,808	125,808	125,808
Mean of Dep. Var.		0.183	0.183	0.113	0.113
Ext-Margin Elasticity		0.395	-0.729	0.474	-0.777

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by zip code appear in parenthesis. For any X, $\Delta Ln(X)$ 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. All specifications include firm fixed effects, sector-period fixed effects, and province-period fixed effects.

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

	Panel A: Exclude zipcodes w/ high auto employment share			Panel B: Exclude zipcodes with at least one sizeable auto maker		
Dependent Variable:	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	1st Stage	2nd Stage	OLS	1st Stage	2nd Stage
Δ Ln(Total Sales)	0.787^a		-4.317^a	0.771^a		-6.016^b
	(0.039)		(1.409)	(0.047)		(2.393)
ΔLn(Vehicles p.c. in Municipality)		0.181^a			0.147^a	
		(0.056)			(0.047)	
Δ Ln(TFP)	0.381^a	0.975^a	5.341^a	0.407^a	0.952^a	6.845^a
	(0.057)	(0.025)	(1.369)	(0.071)	(0.032)	(2.275)
Δ Ln(Avg. Wages)	-0.062	-0.419^a	-2.198^a	-0.124*	-0.400^a	-2.831^a
	(0.049)	(0.030)	(0.608)	(0.067)	(0.037)	(0.992)
F-statistic		14.16			6.98	
Observations	7,178	7,178	7,178	4,613	4,613	4,613

	Panel C: Exclude zipcodes 'neighboring' zipcodes in Panel A			Panel D: Exclude sectors w/ I-O links to automakers		
Dependent Variable:	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)	ΔLn(Exp)	ΔLn(DSales)	ΔLn(Exp)
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	1st Stage	2nd Stage	OLS	1st Stage	2nd Stage
Δ Ln(Total Sales)	0.810^a		-4.851^b	0.825^a		-6.095^b
	(0.040)		(1.888)	(0.046)		(2.469)
ΔLn(Vehicles p.c. in Municipality)		0.170^a			0.136^a	
		(0.056)			(0.047)	
ΔLn(TFP)	0.351^a	0.972^a	5.836^a	0.330^a	0.975^a	7.058^a
	(0.060)	(0.025)	(1.817)	(0.067)	(0.029)	(2.388)
Δ Ln(Avg. Wages)	-0.071	-0.393^a	-2.292^a	-0.118^b	-0.399^a	-2.874^a
	(0.060)	(0.025)	(1.817)	(0.067)	(0.029)	(2.388)
F-statistic		9.38			8.33	
Observations	6,137	6,137	6,137	6,080	6,080	6,080

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include sector and province fixed effects. For any X, ΔLn(X) is the log difference between the average of X in 2009-2013 and its average in 2002-2008. 'Exp' denotes exports, and 'DSales' denotes domestic sales. 'Vehicles p.c.' denotes the stock of vehicles per capita. 'F-statistic' denotes the corresponding statistic for the null hypothesis that the coefficient on the ΔLn(Vehicles p.c. in Municipality) covariate is equal to zero.

Table H.5: Heterogeneous Effects: First Stage

| Sample: | Number of workers is in the interval: | | | | | Low Home Bias | High Home Bias |
| | $(0, 25]$ | $[26, \infty)$ | $[51, \infty)$ | $[101, \infty)$ | $[201, \infty)$ | | |
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1st-Stage Coefficient	0.293^a	0.197^a	0.164^b	0.074	0.043	0.260^a	0.161^a
	(0.086)	(0.049)	(0.066)	(0.084)	(0.117)	(0.071)	(0.058)
F-Statistic	11.73	15.93	6.24	0.78	0.14	13.33	7.67
Observations	2,641	5,376	3,190	1,672	790	4,768	3,249

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include sector and province fixed effects. Vehicles p.c. denotes the stock of vehicles per capita. First-stage coefficient and F-statistic refer to the elasticity of the change in the firm's log domestic sales with respect to the change in log vehicles p.c. in the municipality of location of the firm. For columns 1 to 5, the firm's number of workers is measured as the average across all years the firm appears in the sample.

Table H.6: Heterogeneous Effects: Second Stage

| Sample | Low Labor Elasticity | High Labor Elasticity | Low Materials Elasticity | High Materials Elasticity |
	(1)	(2)	(3)	(4)
2nd-Stage Coefficient	-3.325^c	-1.753^c	-3.169^c	-1.836^c
	(1.872)	(1.013)	(1.792)	(1.040)
1st-Stage F-Stat.	8.80	11.44	6.96	15.26
Observations	3,914	3,914	4,100	3,711

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include sector and province fixed effects. Vehicles p.c. denotes the stock of vehicles per capita. First-stage coefficient and F-statistic refer to the elasticity of the change in the firm's log domestic sales with respect to the change in log vehicles p.c. in the municipality of location of the firm. In columns 1 to 4, we classify firms on the basis of firm-specific labor and materials elasticities computed following the procedure in Bilir and Morales (2018). This table eliminates the specifications in the first two columns of Table 6, as these are not meaningful when ΔLn(Total Sales) is the right-hand-side variable.

Table H.7: Alternative Instruments and Overidentification Tests

Dependent Variable:	ΔLn(Domestic Sales)					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Vehicles p.c. in Province)	0.617^a (0.129)					
ΔLn(Distance-Population Weighted Vehicles p.c. in Other Zip Codes)		0.136^a (0.027)	0.083^a (0.021)			
ΔLn(Vehicles p.c. in Municipality)			0.204^a (0.060)			
Ln(Urban Land Supply Ratio in 1996)				0.018^b (0.009)		
ΔLn(Construction Wage Bill) × 2002 Wage Bill Share in Municipality					0.286^a (0.050)	
ΔLn(Foreign Tourists) × 2002 Foreign Tourists p.c. in Province						0.176^a (0.064)
F-statistic	22.80	25.48	16.97	4.13	32.82	7.55

	ΔLn(Exports)					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔLn(Total Sales)	-1.972^b (0.785)	-3.120^b (1.537)	-2.100^b (0.877)	-2.605 (2.047)	-1.819^b (0.745)	-1.869^a (0.537)
P-value for Sargan Test	0.76	0.23	0.23	0.60	0.56	0.92
Observations	8,018	7,949	7,949	6,940	7,928	8,018

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by province except for columns 4 and 5, in which they are clustered by municipality. 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, and columns 4 and 5 also include province fixed effects.

Table H.8: Confounding Factors

Dependent Variable:	ΔLn(Exports)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔLn(Total Sales)	-2.624^a	-2.347^a	-2.724^a	-2.783^a	-2.143^b	-2.096^b	-2.721^a
	(0.916)	(0.888)	(1.019)	(1.021)	(0.881)	(0.854)	(1.017)
ΔShare of Temp. Workers (firm level)		-0.501^b (0.199)					
ΔShare of Temp. Workers (municipality level)			-0.240 (0.259)				
ΔManufacturing Empl. p.c. (municipality level)				-0.422^a (0.113)			
ΔLn(Financial Costs) (firm level)					-0.057^a (0.020)		
Ln(Financial Costs in Boom) (firm level)						-0.007 (0.020)	
Ln(Financial Costs in Boom) (municipality level)							-0.006 (0.063)
F-Statistic	23.42	21.35	21.00	20.49	20.18	21.23	21.43
Observations	8,018	7,649	7,746	7,748	6,886	6,952	7,743

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

Table H.9: Alternative TFP Measures

Dependent Variable:	ΔLn(Exports)			
	(1) OLS	(2) IV	(3) OLS	(4) IV
ΔLn(Total Sales)	0.785^a	-2.624^a	0.880^a	-1.969^b
	(0.037)	(0.916)	(0.028)	(0.931)
ΔLn(Avg. Wages)	-0.095^b	-1.567^a	-0.293^a	-1.097^a
	(0.047)	(0.408)	(0.063)	(0.288)
ΔLn(TFP Sales)	0.397^a	3.741^a		
	(0.053)	(0.900)		
ΔLn(TFP Value Added)			0.504^a	1.584^a
			(0.056)	(0.365)
F-Statistic		23.42		17.85
Observations	8,018	8,018	8,018	8,018

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality. All specifications include firm-level log average wages and sector and province fixed effects as additional controls.

Table H.10: Placebo Tests of First Stage

Dependent Variable:	ΔLn(Domestic Sales)			
Sample:	Boom firms		Bust firms	
	Within Boom	Boom vs. Bust	Within Bust	Boom vs. Bust
	(1)	(2)	(3)	(4)
ΔLn(Vehicles p.c. in Municipality)	-0.087^c	0.113^b	0.056	0.128^a
	(0.053)	(0.048)	(0.063)	(0.048)
Observations	5,344	5,344	5,245	5,245
F-statistic	2.73	5.40	0.79	7.12

Note: a denotes 1% significance, b denotes 5% significance, c denotes 10% significance. Standard errors clustered by municipality appear in parenthesis. All specifications include firm-level log TFP and log average wages as additional controls. These coefficients are not included to save space. All specifications also include sector and province fixed effects. The sample use to compute the estimates in columns 1 and 2 includes all firms active in at least one year in the subperiod 2002-05 and in at least one year in the subperiod 2006-08. The sample use to compute the estimates in columns 3 and 4 includes all firms active in at least one year in the subperiod 2009-11 and in at least one year in the subperiod 2012-13.

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