Export price adjustments under financial constraints *

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Abstract

Exploiting data on product-destination level transactions of a large panel of Italian firms, we provide new evidence on the effect of financial constraints on price variation across exporters. Controlling for firm characteristics and endogeneity of constraints, constrained exporters charge higher prices than unconstrained firms exporting in the same product-destination market. The positive price difference increases with horizontal differentiation of products, while it reduces in vertically differentiated products, where there is more scope for quality adjustments. The results are consistent with constrained firms exploiting demand rigidities to keep prices up in the attempt to sustain revenues and escape the constraints.

JEL codes: F10, F14, F36, G20, G32, L25

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1 Introduction

Informational asymmetries and imperfect screening in capital or credit markets give raise to situations where financing constraints prevent potentially successful and growth enhancing projects or businesses to be undertaken. There are a few reasons why access to finance plays a particularly important role for firms involved in export activities.[1] Firstly, the need to build ad-hoc distributional networks, to acquire specific information on destination markets, or to customise products, all the way to the mere transportation of goods imply that fixed and variable costs tend to be higher for exporters. Moreover the time lag between production and actual realization of the corresponding revenues is, in general, longer, and international sales contracts are usually more complex, riskier and less enforceable than the domestic ones. Building upon these ideas, an increasing number of empirical papers have recently focused on the effects of financial frictions on firms' exporting activities. These works show that financial constraints reduce firms' ability to enter international markets and the volume of trade, and limit exporters' product scope as well as the number of trade partners.[2]

While financial constraints are found to be critical determinants of trade in many respects, their relationship with export prices is less investigated. This seems particularly unfortunate. First, because it is quite intuitive that constraints can distort firms from optimal price-setting just as they affect optimal extensive or intensive margins. In addition, price adjustments can be seen in a sense as an even more natural response of firms to problems in accessing finance than adjustments along the margins of export. Changing prices is in principle faster and easier than adjusting quantities or changing product/geographical diversification, although of course the strength of competition, demand characteristics, quality and other factors might all influence the scope to maneuver on pricing strategies.

This paper contributes to fill the gap in the literature by providing an exhaustive analysis of the role that limited access to external finance plays in influencing price variation across exporting firms and by uncovering the main factors operating behind this relationship. Understanding the influence of constraints on export pricing is of great relevance not only for the implications on the export performance of single firms, but also for the impact likely induced at the aggregate level. According to the recent international trade literature, indeed, changes in prices play a crucial role in linking trade liberalization to aggregate productivity dynamics. Fiercer import competition induced by the opening up to trade forces to adjust prices and profits, triggering a process of market shares reallocation that leads to sectoral productivity improvements. The recent financial crisis and the evidence on the related contraction of international trade make our investigations particularly relevant.[3]

We exploit an original Italian database obtained by merging a firm-level dataset, including standard balance sheet information, with a transaction-level dataset, recording custom information on

[1]See Manova et al. (2011) among others.

[2]Consistent results are presented in Muuls (2008) for Belgium, in Bellone et al. (2010) for France, in Minetti and Zhu (2011) for a cross-section of Italian firms, and in Li and Yu (2009) and Manova et al. (2011) for Chinese firms. The only contrasting evidence is in Greenaway et al. (2007) for UK, where the probability of entry into exporting is not affected by financing problems, and in Berman and Hricourt (2010), where financing problems do not influence export values in a sample of nine developing countries.

[3]See Amiti and Weinstein (2011), Levchenko et al. (2010), Feenstra et al. (2011) and Chor and Manova (2012) for recent analyses of the role of financial constraints in restricting trade flows.

values and quantities exported for each product and destination. The key advantage of our data is that we have both a proxy of the price charged by each firm for each product-destination transaction, and an informative and reliable firm-level measure of access to credit. The latter is based on a credit rating index issued by an independent agency and available for all the firms in the dataset. Compared to alternative proxies of financial constraints, either based on balance sheets variables or on surveys eliciting whether firms perceived themselves as constrained, credit ratings incorporate the credit markets' view on a particular firm, thus getting close to the actual way investors' decide to provide external finance. The specific rating index that we use, in particular, has been for long embedded in the Italian banking system. It is thus relied upon by Italian banks and tightly linked with the supply and cost of credit. This makes it particularly compelling, given the disproportionate reliance of Italian firms on bank debt, as compared to other sources of finance more heavily exploited in other major economies.

While few other recent works look at the relationship between financial constraints and export pricing, this paper represents, to the best of our knowledge, the only attempt to explicitly control for possible endogeneity bias due to omitted variables or reverse causality. To overcome these potential problems and to reach a proper identification, we adopt an instrumental variable strategy. Following an established practice in the empirical studies on Italy, we use historical information about changes in the Italian banking regulations to identify exogenous restrictions to the local supply of banking services (Guiso et al., 2004; Minetti and Zhu, 2011). Precisely, we exploit exogenous variation in provincial credit supply determined by the progressive removal, during the 1990s, of a series of restrictions to banking services introduced in 1936 by the Bank of Italy.[4]

Our main result is that financial frictions play a relevant role in influencing export pricing. Italian firms facing tighter credit conditions charge higher prices than unconstrained firms exporting an equal product to the same destination. Such positive price-premium holds true even when we control for a set of firm-level characteristics that might influence firms' export prices, and it is robust to a series of sensitivity analyses concerning different estimation methods, different samples and different measures of performance.

We interpret this finding in light of existing theories both within and outside the trade literature. Models of competition in markets with demand rigidities provide natural candidates to explain why financially constrained firms would charge higher prices. Indeed, under capital market imperfections, firms facing difficulties in obtaining credit need to generate extra internal resources to sustain investment or meet current liabilities. Charging higher prices is a way to fulfill this need. Yet, there must be strong enough frictions in the product market allowing to increase prices without loosing too much of their demand, at least in the short run. In the micro-trade literature, Fan et al. (2012) propose another possible explanation within a monopolistic competition model à la Melitz (2003), extended to include quality, credit constraints and marketing costs. There, the ultimate effect of financial constraints on export prices results from two opposing effects. First, a quality adjustment effect: constrained firms cannot afford the costs of quality and thus sell lower quality goods at lower prices. Second, a price distortion effect that, similarly to the mechanism described above, comes from the ability to charge higher prices exploiting demand rigidities. Finally, there is a third mechanism, consistent with both

[4]See Guiso et al. (2006); Herrera and Minetti (2007); Alessandrini et al. (2010) for other applications in the empirical literature on Italy.

types of models, through which financial frictions might influence export prices: a simple marginal costs effect, according to which constrained firms set higher export prices because of higher unit costs.

We find that the positive price difference between constrained and unconstrained firms survives after controlling for the marginal costs channel. Further, the paper concludes providing novel evidence that both price distortion related to demand rigidities and quality adjustment mechanisms play a role. Taking advantage of the high disaggregation of our dataset, we show that the positive price premium of constrained firms is larger in more horizontally differentiated products, where one expects that there is more room to leverage on price rigidities, while it is offset by quality effects in vertically differentiated products, where there is more scope for quality adjustment.

Within the vast empirical literature of firm heterogeneity in international trade, this article more directly relates to the only two works that analyze the impact of financial frictions on pricing strategies. Manova et al. (2011), using Chinese custom data, find two contrasting results: while MNC affiliates set lower export prices in financially vulnerable sectors, joint ventures have higher unit values in the same industries. Fan et al. (2012), using the same data, provide evidence that firms in sectors with high external finance dependence set, on average, lower prices. These papers share two key limitations. They rely on industry-level rather than on firm-level measures of access to finance to identify the effect of financial constraints on firm performance, and do not control for possible endogeneity of financial constraints. Our analysis overcomes both problems. Additionally, our work also relates to the growing empirical literature documenting the systematic variation in export prices across firms, products and trade partners (Bastos and Silva, 2010; Manova and Zhang, 2012a; Harrigan et al., 2011; Fan et al., 2012). These studies directly link export prices to firm characteristics including productivity, size, capital intensity, and the skill composition of workers. By contrast, our study reveals that firm financial conditions represent a further crucial determinant of product-destination export prices.

The remainder of the article is organised as follows. Section 2 presents the data set, our proxy of financing constraints, and basic descriptive evidence. Section 3 presents the empirical methodology and the main empirical findings. Section 4 discusses the results in light of the different hypotheses proposed by the theoretical literature, and provides evidence on marginal cost, product differentiation and quality channels. Section 5 then concludes.

2 Data and descriptive analysis

In this Section we present the data and provide descriptive evidence on the relationship between our proxy of financial constraints (FCs) and the main variables.

The Data

The analysis combines three sources of data: the Italian Foreign Trade Statistics (COE) and the Italian Register of Active Firms (ASIA), both collected by the Italian Statistical Office (ISTAT), and a firm level accounting dataset (CEBI-CERVED), which is available through ISTAT but collected by

Centrale dei Bilanci (CEBI, the Italian Company Account Data Service).[5]

The COE dataset is the official source for trade flows of Italy. It records separately the value (in Euros) and the quantity (in kilos) involved in each export and import cross-border transaction performed by a firm, thus allowing to compute export and import prices (unit values). As usual with custom-level data, we compute unit values as the ratio between the value and the physical quantity of each export transaction. Because the reported value of exports excludes the cost of insurance and freight, the unit price of exports is a free-on-board (f.o.b) price.[6] Traded products are classified at the six digit level of the Harmonised System (HS6). The data available to the present study cover the period 2000-2003, for a total of $5,329$ product categories exported in 236 different destination countries.

The ASIA register covers the universe of Italian firms active in the same time span, irrespective of their export status. It reports annual figures on number of employees, sector of main activity, and information about geographical location of the firms (municipality of principal activity or legal address).

The CEBI-CERVED dataset collects annual administrative reports for all Italian *limited liability* firms. The long term institutional role of CEBI ensures high data quality, limiting measurement error.[7] The annual reports contain information on financial and non financial variables. The key variables for this study include employment, domestic sales, age, total assets, gross operating margin and a firm-level credit rating index.[8]

We merge these three data sources and obtain a dataset covering the entire population of Italian limited firms (exporters and non exporters). The major advantage of matched firm-trade data is that they enable us to directly relate export prices to firm attributes. The main limitation of the sample rests in a mild over-representation of bigger and more productive firms.[9] We focus on firms operating in manufacturing. Further, since the short time span available (2000-2003) and the inclusion of lagged variables place limits to the exploitation of time variation, our main analysis considers time-series averages of the relevant variables. This avoids an arbitrary choice of a single year, and allows to account for possible individual year shocks. The final sample includes a total of $114,866$ firms.

Columns 1 and 2 of Table 1 presents descriptive statistics about the sample characteristics in terms of key variables employed in the following regression analysis. These include: size measured by the number of employees, age computed by the year of foundation, a TFP measure (in logs) obtained via the IV-GMM modified Levinsohn-Petrin estimator proposed in Wooldridge (2009), and two variables that interact with external credit constraints in determining the financial status of a firm: total assets to proxy for availability of collateral, and gross operating margins as a measure of internally generated

[5]The datasets have been made available for work after careful screening to avoid disclosure of individual information. The data were accessed at the ISTAT facilities in Rome.

[6]ISTAT collects data on exports based on transactions. The European Union sets a common rules for data collection across countries, but leaves some flexibility to member states. A detailed description of requirements for inclusion in Italian export data is provided in Appendix.

[7]Centrale dei Bilanci (CEBI) was founded as a joint agency of the Bank of Italy and the Italian Banking Association in the early 1980s to assist in supervising risk exposure of the Italian banking system. Today part of CERVED, the leading group in business information services in Italy, during the sample period CEBI was a private company owned by major Italian banks which exploited its services in gathering and sharing information about firms.

[8]We also exploit value added, cost of materials and tangible assets to estimate a TFP measure, see below.

[9]Further details on the data sources and their coverage are reported in Tables A1 and A2 in the Appendix.

Table 1: DESCRIPTIVE STATISTICS

	Coverage of Italian Manufacturing (1)	Our sample - Averages (2)	Difference between FC and non-FC firms (3)
Panel A - All firms			
Number of firms	101,546		
Percentage of firms[a]	0.21		
Number of employees		28.668	-0.840*** (0.013)
Age		13.959	-0.589*** (0.011)
log TFP		2.562	-0.447*** (0.013)
Total Assets		6,312.1	-0.982*** (0.017)
Gross operating margin		596.4	-2.590*** (0.022)
Panel B - Exporters			
Number of firms	48,347		
Percentage of firms[a]	0.59		
Share of export value[a]	0.84		
Number of employees		50.203	-0.751*** (0.028)
Age		16.987	-0.694*** (0.022)
log TFP		2.734	-0.463*** (0.022)
Total Assets		11,768.1	-0.733*** (0.033)
Gross operating margin		1,138.4	-2.540*** (0.051)

Notes: Figures computed on time-series averages over 2001-2003. [a] Figures on coverage refer to 2003, similar results in the other years.

Panel A - Column 1: coverage of the sample with respect to aggregate Italian manufacturing. Column 2: averages of number of employees, age, Wooldridge (2009) modified Levinsohn-Petrin TFP, total assets, and gross operating margins, all computed across all firms. Column 3: difference in means between constrained and unconstrained firms in the entire sample: log-OLS regressions controlling for 3-digit industry.

Panel B - Column 1: coverage of the sample in terms of number of exporters and export value with respect to aggregate Italian manufacturing. Column 2: averages of number of employees, age, Wooldridge (2009) modified Levinsohn-Petrin TFP, total assets, and gross operating margins, all computed across exporting firms in the sample. Column 3: difference in means between constrained vs. unconstrained exporters: log-OLS regressions controlling for 3-digit industry.

Robust standard error in parenthesis. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

resources.[10]

Compared to aggregate Italian manufacturing (column 1), we cover 21% of firms, about 59% of all manufacturers that do export, and 84% of the total value of exports performed by Italian manufacturing firms.[11] Comparing all firms vs. exporting firms (column 2, Panel A vs. Panel B), we confirm the stylised facts that exporters are on average bigger, older and more productive. Also, they have a stronger financial side, with more assets and more internal resources.

Financial constraints

We build a measure of financial constraints based on the credit rating index issued yearly by CEBI, available through the CEBI-CERVED dataset. The score results from a multivariate analysis that summarizes a wide range of financial and non-financial characteristics of the firm.[12] The original index ranks firms in 9 categories of creditworthiness: 1-high reliability, 2-reliability, 3-ample solvency, 4-solvency, 5-vulnerability, 6-high vulnerability, 7-risk, 8-high risk, and 9-extremely high risk.[13] In keeping with the binary categorization traditionally employed in the literature, we build a financial constraints dummy (FC) that equals 1 if a firm is rated 8 or 9, and 0 otherwise. The index is updated at the end of each year. It is therefore the rating in $t-1$ that is relevant for credit suppliers' current decisions on credit provision.

The use of a credit score to proxy for firms' access to finance has both advantages and limitations, as many other alternatives explored in the literature. The main strength is that ratings, in general, closely influence "the opinion [of the markets] on the future obligor's capacity to meet its financial obligations" (see Crouhy et al., 2001), thus capturing the actual propensity of potential investor to grant credit. Our score, in particular, is perceived as an official rating and it is thus used by Italian banks as a benchmark for internal procedures for the evaluation of potential borrowers. This is crucial in the Italian case, where firms strongly rely on bank credit due to underdevelopment of bond and stock markets. In fact, although the Italian banking system is comparatively small with respect to the real economy (2.7 times the GDP compared to, for instance, 4.2 times the GDP in France), banks play a prominent role in the financing of firms in Italy. Indeed almost 70% of the financial debts of non-financial corporations is made up by bank loans, while the same share is only 37% in France and 55% in Germany (see Figure A1 in the Appendix). By contrast, Italian capital and bond markets are quite small compared to other major countries. The stock market capitalization of Italian non financial corporations is less that 20% of GDP, compared with 75% in France and 45% in Germany (Figure A2 in the Appendix). Bond financing of Italian non financial corporations amounts to less than 8% of firms' total financial debt (Figure A3 in the Appendix). In this context, the rating score captures crucial aspects of credit supply and it works well as a proxy for what banks do.

[10]In applying the Wooldridge (2009) TFP estimator, we take value added as a proxy for output, employees and gross tangible assets to proxy labour and capital inputs, and cost of material inputs as a proxy for intermediate inputs.

[11]These numbers refer to year 2003 but similar values are observed for the other years.

[12]While the method to construct the rating index is proprietary information of CEBI-CERVED, it is known that information on firms' international activities does not enter the score.

[13]These definitions are valid over the sample time period. Changes in the definition and the number of score classes occurred afterwards, following subsequent changes in CEBI ownership and updates in rating procedures.

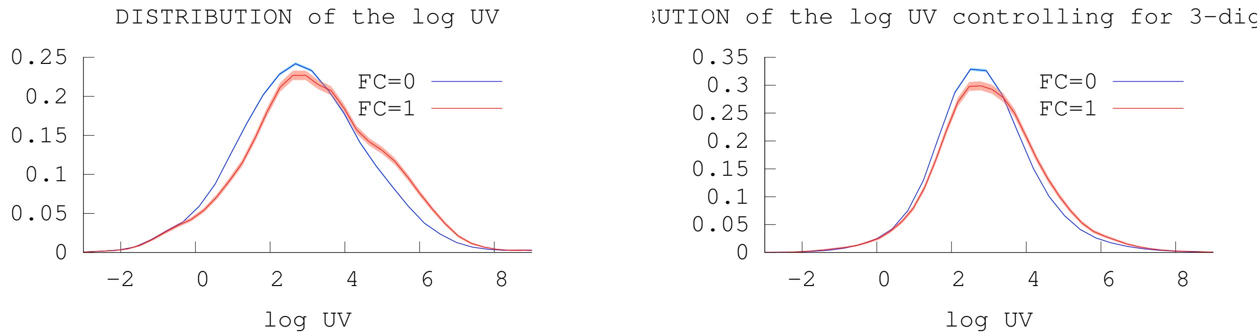

Figure 1: EMPIRICAL DENSITY OF (log) UNIT VALUES.

Notes: Kernel estimate of the empirical density of the (log) unit value of export transactions performed by constrained (red line) and unconstrained (blue line) exporters, not controlling (left-panel) and controlling (right panel) for 3-digit industry. The kernel function is Epanenchnikov and the bandwidth is set according the standard heuristics in Silverman (1986). A conservative confidence band (confidence level higher than 99%) is also reported. Stochastic dominance of the density of constrained firms is verified via the (Fligner and Policello, 1981) test. The value of the test statistics is 60.54 (left-panel) and 40.10 (right-panel), with associated p-values lower than 10^{-6}.

Our credit rating enjoys other features that make it particularly suitable to measure financial constraints. Firstly, it is available for all firms in the sample, while scores from international credit agencies are biased toward a smaller subset of Italian firms. Secondly, as an at least indirect proof of its actual relevance in banks' lending decisions, there is a tight link with the availability and cost of external finance: Pistaferri et al. (2013) show that it is unlikely that a firm with poor rating can receive any credit, while Panetta et al. (2009) provide clear evidence that bad ratings have a clear association with higher cost of credit. Finally, Bottazzi et al. (2008, 2013) show that an important fraction of highly productive, highly profitable and fast growing firms receive poor scores. Hence our index is not simply a summary measure of firm performance, but it actually captures a more complex set of information that a bank would consider when lending to firms.

In column 3 of Table 1 we present basic correlation between our measure of financial constraints and some key firm characteristics. We report differences in mean between constrained and unconstrained firms by running an OLS regression of firm attributes (in logs) on the FC dummy, including 3-digit industry fixed effects to get rid of sector-specific patterns on the production side. Looking at the entire sample (in Panel A) we confirm common findings about constrained firms: FC firms tend to be smaller, younger, less productive, and suffer, on average, from a relative weaker financial structure in terms of less assets and less internally generated resources. Once we condition on being exporters (in Panel B), the results do not change.

Figure 1 starts looking at the relationship between constraints and prices, by comparing the empirical distributions of (log) unit values of export transactions performed by constrained (red line) and unconstrained firms (blue line). In the left panel we do not control for 3-digit industry effects, in the right panel we do via computing deviations from industry averages. Two comments are in order. First, the observed density for FC firms is shifted rightward suggesting that these firms tend to charge higher unit values. The visual impression is validated by means of a Fligner-Pollicello test for stochastic dominance, a non parametric test robust to various forms of asymmetry and heteroskedasticity in the samples (Fligner and Policello, 1981). The null hypothesis is that the probability that a FC firm displays a higher (log) unit value than a non-FC firm is higher than 1/2. The hypothesis cannot be rejected in our sample, irrespective of whether we control or not for 3-digit industries. Second, the support of the densities is wide: observed (log) unit values range approximately from -2 to 8. This is not a surprising fact, as we are indeed mixing together unit values of, e.g., 'pasta' (with a median 0.62 €/Kg) and 'precious metals' (with a median 17,200€/Kg), even when we do control for 3-digit industry averages. This strongly supports that the identification of the relation between financial constraints and unit values cannot avoid to control for detailed product and country characteristics, as we indeed seek to do in the following via product-destination fixed effects.

3 Empirical analysis

In this Section we move to regression analysis. We introduce the baseline empirical model and the identification strategy, and then present our main findings.

Empirical model

The relationship between export prices of Italian manufacturing firms and their FC status is studied through the following baseline regression

$$\ln \mathrm{EUV}_{fpc,t} = \gamma \mathrm{FC}_{f,t-1} + \mathbf{X}'_{f,t-1}\boldsymbol{\beta} + \mu_{pc} + \epsilon_{fpc,t} \quad , \tag{1}$$

where EUV_{fpc} is the (f.o.b) unit value of product p exported to country c by firm f at time t, and the regressor of primary interest is the dummy $\mathrm{FC}_{f,t-1}$, which equals 1 if firm f is financially constrained at time $t-1$ and 0 otherwise.[14] $\mathbf{X}$ is a set of firm-level controls, all measured at $t-1$. It includes all the variables already presented in the descriptive analysis above. These are productivity measured via the Wooldridge (2009) modified Levinsohn-Petrin estimator of TFP (in logs), size measured by the (log) number of employees, (log) age computed by the year of foundation, and two more financial-side: (log of) gross operating margins as a measure of internal financial resources, and (log of) total assets to proxy for overall availability of collateral. The controls in $\mathbf{X}$ also include geographical dummies (North, Center, South), accounting for well known differences in the level of development and other characteristics across Italian regions, as well as a full set of 3-digit industry dummies.[15] We use the lagged value of firm characteristics to ensure that the estimated coefficients are not contaminated by possible feedback effects of export prices on productivity and other firm attributes. The error term includes a product-destination fixed effect, μ_{pc}, and a standard random component ϵ_{fpc}. Product-destination fixed effects control for all factors common across firms active within the same product-destination pair, including fixed export costs specific to the product variety and the partner country.

[14] Recall that t and $t-1$ in equation 1 represent time averages across 2001-2003 and 2000-2002, respectively.

[15] Nominal variables are deflated with appropriate sectoral price indexes collected by ISTAT. Complete deflator series are available only at the 2-digit level. We therefore perform deflation at this level of aggregation. The base year is 2000.

Table 2: FINANCIAL CONSTRAINTS AND (log) UNIT VALUE: OLS-FE estimates

	(1) No controls	(2) Baseline	(3) Empl>1	(4) Core Products-Countries
Financially constrained firms dummy (FC)	0.159*** (0.005)	0.209*** (0.005)	0.207*** (0.006)	0.143*** (0.007)
log TFP		0.254*** (0.002)	0.259*** (0.002)	0.246*** (0.003)
log number of employees		0.168*** (0.002)	0.174*** (0.002)	0.152*** (0.002)
log age		0.023*** (0.001)	0.024*** (0.001)	0.029*** (0.002)
log total assets		-0.120*** (0.001)	-0.123*** (0.001)	-0.134*** (0.002)
log gross operating margin		-0.043*** (0.001)	-0.044*** (0.001)	-0.044*** (0.001)
North dummy		0.134*** (0.004)	0.132*** (0.004)	0.172*** (0.005)
Center dummy		0.131*** (0.004)	0.130*** (0.004)	0.138*** (0.006)
Number of observations	2,454,168	2,454,168	2,439,626	1,050,125
Product-Country FE	Yes	Yes	Yes	Yes
3 dgt sectoral dummies	Yes	Yes	Yes	Yes

Notes: The dependent variable is the (log) export unit value at the product-country level. Columns 1 and 2: regressions estimated on the whole sample of exporters. Column 3: regression estimated on the whole sample of exporters after removing firms with 1 employee. Column 4: regression estimated on the whole sample of exporters after removing transactions in products and destinations whose share in a firm total export value is below 1%. All the regressions include a constant, 3-digit industry dummies and HS6 product-country pair fixed effects (170,664 categories in columns 1-2; 170,187 in column 3; 98,944 in column 4).
Robust standard errors in parenthesis, clustered at product-country level. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

They therefore also account for product-specific characteristics that are invariant across manufacturers and trade partners, as well as destination-specific economic determinants of prices that affect all products and firms selling there, such as consumer income, bilateral distance, transportation costs, market toughness.

Basic correlation is captured by OLS estimates of Equation 1. Given product-destination fixed effects, the coefficient of interest (γ) is identified purely from the variation of export prices and firm characteristics across firms within the same product-destination market. Results are reported in Table 2 with standard errors clustered at product-country level. In column 1, we suppress firm level controls. We confirm the positive correlation of financial constraints with unit values suggested by the distributional analysis above. The estimated coefficient tells that the price charged by constrained firms is 17.2% (0.159 log points) higher than the price set by unconstrained firms exporting the same product to the same destination market.

In column 2, we introduce firm specific controls. The estimate of γ remains positive and the price gap increases to 23.2% (0.209 log points). This implies that failing to control for other firm attributes associates with a downward bias in the γ coefficient. For instance, since it is likely (as found in previous studies) that less productive or smaller firms tend to fix lower prices and are also more likely to be financially constrained, excluding productivity or size from the regression understates the impact of FCs on export prices. With respect to the specification without controls the observed effect is 31% higher: about a third of the price gap is associated with the differences between constrained and unconstrained firms in productivity, size, age, total assets and internal resources.

The estimated coefficient on (log) TFP suggests that a 100 percent increase in productivity is associated with a 28.9% (0.254 log points) increase in export prices. The result corroborates previous findings and it is consistent with the endogenous quality theory: more productive firms choose more expensive inputs to produce higher quality goods for which they charge higher prices (Bastos and Silva, 2010; Ge et al., 2013). The coefficient on (log) employees suggests that the price elasticity of firm size is also positive and significant: everything else equal, bigger firms charge higher prices for a product exported in a certain destination (Kugler and Verhoogen, 2012). The magnitude of the association is somewhat smaller than the one observed for productivity: a 100 percent increase in firm size translates into a 18.3% increase in export price. The two specific financial controls, gross operating margins and total assets, both turn out significant and with a consistent negative sign: the greater the availability of internal resources or collateral (and thus the lower the likelihood to be constrained), the lower the unit values charged for exported products. The associated percentage reduction in unit values is 12.7% and 4.4%, respectively. We also observe that firms located in the center-north of Italy tend to set higher unit values. This is expected, however: given the geographical disparities in Italy, center-north located firms are likely to be more innovate and to produce higher quality products. Finally, firm age is the only firm characteristic for which we do not observe statistically significant results.

In column 3 and column 4 we perform two robustness checks considering two different subsamples. First, we remove the firms reporting only one employee. These firms represent self-employment, headquarters, or ultimate parent companies in groups. Their export dynamics are thus very specific. The estimated coefficient on the FC dummy and on the other firm attributes remain unchanged (column 3) compared to the full-sample results. Second, we drop marginal products and destinations, defined here as those involving less than 1% of the overall exports of each firm. Removing such irrelevant transactions might make the identification cleaner, as indeed studies on multi-products firms find that products closer to the core firm competencies sell for higher prices than non-core products (Manova and Zhang, 2012b; Eckel et al., 2011). The main conclusion of positive price premium for FC firms (column 4) is robust to this additional control. The estimated impact of FCs is smaller (0.143 log points, i.e. 15.4%), suggesting that constrained firms are more focused on core products and destinations than unconstrained firms.

Identification strategy

Comparing OLS regressions with and without firm-level controls points at a standard omitted variable bias whose direction is quite intuitive: since banks are more likely to provide credit to larger, more productive and less financially vulnerable firms, omitted variables associate with a downward bias in γ. Introducing firm level controls helps in mitigating this problem, but one can never be sure to completely eliminate the bias, especially because firm fixed effects are not viable given the time-series average structure of the data. Moreover, there might be a second form of endogeneity which we

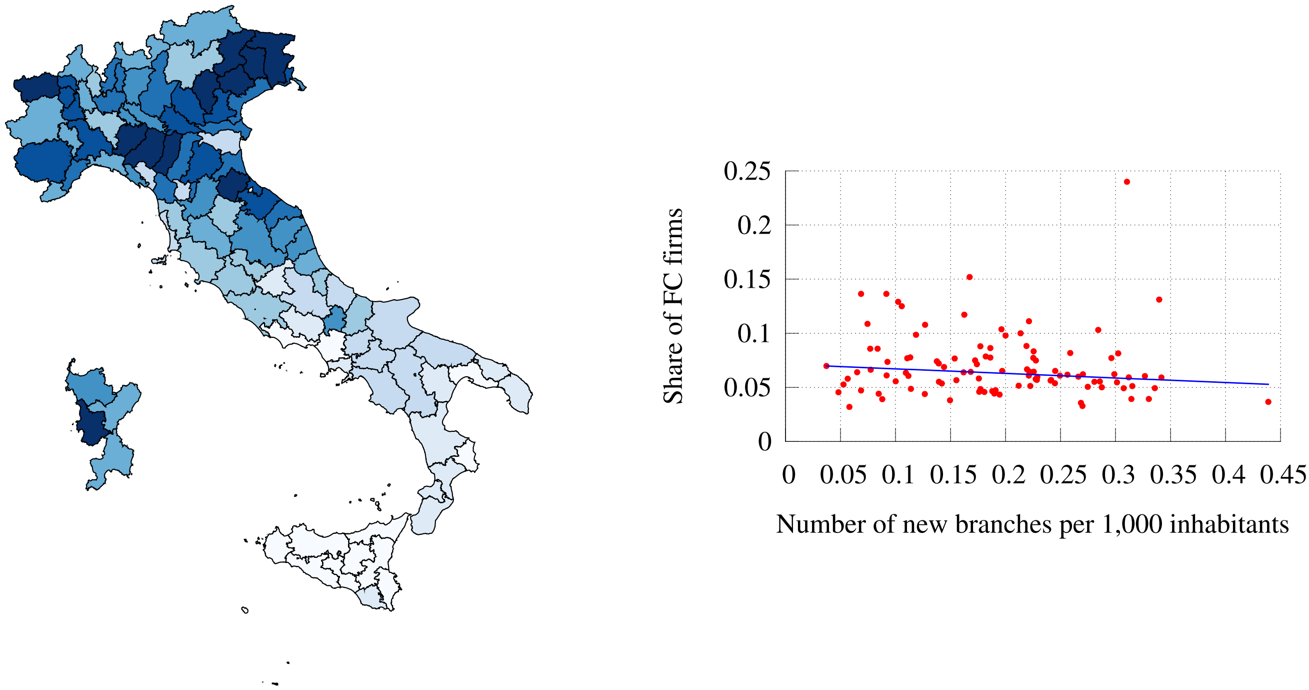

Figure 2: GEOGRAPHICAL DISTRIBUTION OF AVERAGE NUMBER OF NEW BANK BRANCHES OPENED YEARLY DURING 1990-99 (LEFT PANEL). RELATION BETWEEN NEW BANK BRANCHES OPENED IN 1990-99 AND SHARE OF FC FIRMS, BY PROVINCE (RIGHT PANEL).

Notes: Plots and fit are based on 103 provincial observations. In the left panel, the darker the province map and the higher the number of firms located in that province. In the right panel, the linear fit is a Least Absolute Deviation regression: estimated parameters are 0.052***(0.004) and -0.047***(0.016) for the intercept and the slope, respectively. Robust standard errors, ***: significant at the 1% level.

need to tackle. Indeed, it could be the case that export pricing decisions and the FC status are jointly determined. We use an instrumental variable approach to achieve identification of exogenous variation in firm level access to credit. Good instruments must be exogenous, help to explain the endogenous variable and satisfy the exclusion restriction. In what follows we define our instrument and we explain why we believe it is fairly exogenous. We also provide evidence of its positive correlation with the endogenous variable FC.

In the absence of firm level variables allowing to identify exogenous variation in FC, we exploit the exogenous shock to the geographical variation in credit supply caused by the progressive removal, during the 1990s, of a series of local (provincial level) restrictions to banking services introduced as early as 1936 by the Bank of Italy under the fascist regime. This strategy, originally developed in Guiso et al. (2004), has been recently applied by Minetti and Zhu (2011) to study how FCs affect firm-level export margins of a sample of Italian firms, while here we extend to transaction-level export prices.

As explained in detail in Guiso et al. (2004), the geographical distribution of banks and bank branches across Italian provinces in 1936 came about as the response to the norms enacted by the regulatory authority and it was essentially unrelated with the structural characteristics and the level of development of the provinces themselves. The subsequent removal of the regulation during the 1990s freed up banks' possibility to open new affiliates, with differentiated impact across provinces.

The impact of deregulation varies also in relation to the different types of banks active at the local level, as indeed the 1936 was imposing weaker restrictions to saving banks while more stringent limits to cooperative banks. This exogenous variation is expected to directly affect the probability to be financially constrained. Indeed, the higher the number of new affiliates created in a province, the higher is the availability and the lower the cost of credit in that geographical area. In turn, easy access to bank loans and lower lending costs influence directly the rating index which, as stressed before, is tightly linked with the availability and the cost of external finance. By contrast, the exogenous variation in the supply of credit is not expected to directly impact neither on firm export behavior nor on unobserved firm characteristics that determine export behavior.

In this spirit, we instrument the FC dummy with the 1990-1998 average number of branches (per 1,000 inhabitants) created annually in each province by banks. Left panel in Figure 2 shows the intensity of the phenomenon captured by our IV, by province. The provinces with the greater number of newly created bank branches are those in the center-north of Italy. Still, the instrument shows a great deal of variability, even within sub-areas.[16] A crucial piece of information on the validity of the instrument is provided in the right panel of Figure 2, showing that the instrument is highly correlated with the share of financially constrained firms in each province. A Least Absolute Deviation estimate of the slope of the relationship gives a coefficient of -0.047 (standard error 0.016), providing further support to the instrument.

Once we allow for endogeneity of the FC dummy, Equation (1) becomes a standard dummy endogenous variable model. Following Wooldridge (2010), we estimate the model via a two-step procedure: (i) we estimate by maximum likelihood the binary response model $P(\text{FC}_f = 1|\mathbf{X},\text{Z})$ where **X** is the above set of firm level controls and Z the instrument, and obtain the associated fitted probabilities $\hat{P}$; next, (ii) we estimate equation (1) by 2SLS-IV using the fitted probabilities $\hat{P}$ as instrument. There are several nice features of this IV estimator: it is robust to mis-specification of the probit model, it is more efficient than directly including the number of branches opened in 1990-99 as an instrument into an IV procedure and, finally, it does not require to adjust the 2SLS-IV standard errors.[17]

2SLS-IV results

In the first step we build $\hat{P}$ via maximum likelihood estimation of the following probit

$$\text{P}(\text{FC}_f{=}1 \mid \mathbf{X},\text{Z}) = \Phi\left(\delta_1 \text{Z} + \mathbf{X}'_f \boldsymbol{\beta} + \epsilon_f\right) \quad , \tag{2}$$

where the probability to be in the FC group is regressed on the instrument Z, i.e. the average number of new branches opened over the period 1990-1998 in the province wherein a firm is located, and on the set **X** of (lagged) firm level controls described above.

[16] With this respect, notice that the provinces with an higher number of newly created branches also correspond to provinces where the vast majority of firm is located. In particular: 70% of firms in the sample are located in the North, 18% in the Center and 12% in the South. Robustness checks presented in the following show that our main result is not driven by a simple spatial effect.

[17] Standard weak instrument diagnostics are known to fail in this context (Nichols, 2007). See the Appendix for a number of further validation of the good properties of the instrument $\hat{P}$.

Table 3: PROBABILITY OF FINANCIAL CONSTRAINTS

	(1) Baseline		(2) Empl>1	
Panel A - Exporters				
Number of new branches	-0.589**	(0.273)	-0.689**	(0.275)
log TFP	-0.055***	(0.022)	-0.099***	(0.026)
log number of employees	-0.055***	(0.017)	-0.053**	(0.022)
log age	-0.350***	(0.018)	-0.362***	(0.020)
log total assets	0.208***	(0.018)	0.202***	(0.021)
log gross operating margin	-0.328***	(0.012)	-0.320***	(0.012)
North dummy	0.312***	(0.066)	0.351***	(0.067)
Center dummy	0.280***	(0.069)	0.309***	(0.069)
Panel B - Goodness of fit				
Number of observations	48,347		46,959	
Adjusted R^2	0.254		0.260	
Brier score	0.032		0.037	
AUC score	0.871		0.875	

Notes: Panel A - Probit estimates, the dependent variable is Pr(FC=1). Regressions include a constant and 3-digit industry dummies. Column 1: estimates on the whole sample of exporters. Column 2: estimates on the whole sample of exporters after removing those firms with 1 employee.
Panel B - Goodness of fit statistics. The Brier score is computed as $BS = (1/N)\sum(\hat{FC} - FC)$, where $\hat{FC}$ is status predicted by the model and FC the actual status: the closer to zero the better the fit. The AUC score measures the area under the ROC (Receiver Operating Characteristics) curve: the closer to 1 the better the predictive power.
Robust standard errors in parenthesis, clustered at provincial level. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Table 3 reports the results. The coefficient on the instrument, δ_1, is significant and negative: the more branches were opened in a province, the lower the probability for a firm in that province to face binding financial constraints. Coefficient estimates on the controls tell that more productive, larger, older firms generating higher internal resources are less exposed to credit problems. This is in line with economic intuition. Panel B of Table 3 reports different goodness of fit statistics for the probit model: they do not suggest any specific pathology. The results are consistent also for the alternative sample, that we are going to use as robustness check in the following, obtained by removing the exporters with only one employee (cfr. column 2).[18]

Next, in the second step, we estimate equation (1) via 2SLS-IV taking $\hat{P}$ as the instrument for the FC dummy. Table 4 shows the results. Panel A reports the second stage IV estimates, while Panel B documents the first stage. In column 1 we consider the baseline model over the whole sample of exporters. The instrument works well: as seen above it has explanatory power, and we observe here

[18] Additional exercises supporting the probit goodness of fit are presented in Appendix.

Table 4: FINANCIAL CONSTRAINTS AND (log) UNIT VALUE: IV-FE estimates

	(1) Baseline	(2) Empl>1	(3) Core Products-Countries	(4) No MNCs	(5) Euro Countries
Panel A - Second Stage on (log) unit value					
Financially constrained firms dummy (FC)	0.223*** (0.035)	0.182*** (0.035)	0.740*** (0.042)	0.180*** (0.038)	0.282*** (0.035)
log TFP	0.254*** (0.002)	0.259*** (0.002)	0.253*** (0.003)	0.255*** (0.002)	0.267*** (0.002)
log number of employees	0.167*** (0.002)	0.174*** (0.002)	0.156*** (0.002)	0.176*** (0.002)	0.180*** (0.002)
log age	0.023*** (0.001)	0.023*** (0.002)	0.043*** (0.002)	0.016*** (0.001)	0.024*** (0.001)
log total assets	-0.120*** (0.002)	-0.122*** (0.002)	-0.156*** (0.003)	-0.124*** (0.002)	-0.128*** (0.002)
log gross operating margin	-0.043*** (0.001)	-0.045*** (0.001)	-0.023*** (0.002)	-0.045*** (0.002)	-0.043*** (0.001)
North dummy	0.134*** (0.004)	0.132*** (0.004)	0.162*** (0.006)	0.144*** (0.004)	0.123*** (0.004)
Center dummy	0.131*** (0.004)	0.131*** (0.004)	0.124*** (0.006)	0.137*** (0.005)	0.125*** (0.005)
Panel B - First Stage on FC					
Instrument ($\hat{P}$)	0.559*** (0.006)	0.562*** (0.006)	0.648*** (0.010)	0.528*** (0.007)	0.610*** (0.007)
log TFP	-0.021*** (0.0004)	-0.019*** (0.0004)	-0.021*** (0.001)	-0.020*** (0.0004)	-0.016*** (0.0004)
log number of employees	-0.012*** (0.0003)	-0.010*** (0.0003)	-0.010*** (0.0004)	-0.009*** (0.0003)	-0.009*** (0.0003)
log age	-0.009*** (0.001)	-0.010*** (0.0002)	-0.010*** (0.0003)	-0.012*** (0.0002)	-0.007*** (0.0002)
log total assets	0.020*** (0.0004)	0.020*** (0.0003)	0.013*** (0.0004)	0.019*** (0.0003)	0.019*** (0.0003)
log gross operating margin	-0.005** (0.0003)	-0.005*** (0.0003)	-0.001*** (0.0005)	-0.008*** (0.0003)	-0.006*** (0.0003)
North dummy	0.008*** (0.001)	0.007*** (0.001)	0.007*** (0.001)	0.011*** (0.001)	0.008*** (0.001)
Center dummy	0.027*** (0.001)	0.025*** (0.001)	0.016*** (0.001)	0.020*** (0.001)	0.028*** (0.001)
Number of observations	2,454,168	2,439,626	1,050,125	2,229,799	2,294,173
Product-Country FE	Yes	Yes	Yes	Yes	Yes
3 dgt sectoral dummies	Yes	Yes	Yes	Yes	Yes

Notes: Panel A - second stage results: the dependent variable is (log) export unit value at the product-destination level. Panel B - first stage results: the dependent variable is the FC dummy. Column 1: regression estimated on the whole sample of exporters. Column 2: firms with 1 employee are excluded. Column 3: transactions in products and destinations representing less than 1% of a firm's total export value are dropped. Column 4: MNCs are excluded. Column 5: only transactions toward destinations adopting the EURO currency are included. All the regressions include a constant, 3-digit industry dummies and HS6 product-country pair fixed effects. Robust standard errors in parenthesis, clustered at product-country level. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

that it is statistically significant in the first stage. In the main regression, the estimated coefficient on the FC dummy is 0.223. This implies that financially constrained firms fix on average 25% larger unit values for the same product exported to the same destination country. The elasticities of size and productivity are positive and strongly significant, and about of the same magnitude as the effect of constraints. The coefficient on age is also positive and statistically significant, although the effect is smaller in magnitude (0.023 log-points, i.e. about 2%). Conversely, availability of collateral and availability of internal resources associate with lower prices. Finally, firms located in the center-north of Italy tend to show higher unit values.

It is instructive to compare the 2SLS-IV results with the above OLS estimates with product-country fixed effects. Without instrumenting (see column 2 of Table 2), the coefficient on the FC dummy is positive and significant, but smaller than the 2SLS-IV coefficient (0.209 vs. 0.223). There are several explanations for this downward bias, all related to possible correlation between the regressors and the endogenous component of the FC proxy. First, OLS-FE estimates might be attenuated if firms with a less stringent financial constraint are also those with higher market power and higher markup. Second, firms that are financially constrained might tend to be less productive on average, and at the same time to produce lower quality goods for which they can charge lower prices. If our control for productivity is not perfect, this might contribute to the downward bias of the OLS-FE.

The positive relationship between financial constraints and export prices is confirmed by a series of robustness checks. These are reported in columns 2-5 of Table 4.[19] First, FCs associate with higher export prices even if we exclude firms with one employee (column 2). The point estimate of the FC dummy reduces (0.182), suggesting that FC firms are more numerous within the excluded sample of firms with 1 employees, possibly in the form of self-employing entrepreneurs. Nevertheless, the effect is statistically equal to the baseline result in column 1 within a 1-standard error confidence band. Second, we exclude marginal products and destinations, that is those accounting for less than 1% of the overall export values of a firm (column 3). The positive price premium for FC firms is preserved, but the point estimate on the FC dummy coefficient is now larger (0.740). Since we expect exports of less important products and destinations to be cheaper than exports in core products or countries, the reduction in the FC coefficient suggests that, in line with OLS-FE estimates in Table 2 above, constrained firms are more focused on core products and destinations than unconstrained firms. Third, in order to account for the particular behavior of multinational companies, which are likely to set higher prices (cfr. Ge et al., 2013) than domestic firms, we drop them from our regression. Column 4 shows that our baseline findings are not affected by the behaviour of these firms. Fourth, in order to eliminate possible confounding factors related to exchange rate dynamics, we re-estimate the model considering only the export transactions to partner countries that use the EURO currency over the entire sample period (column 5).[20] The point estimate on the FC coefficient (0.282) is statistically equal to the baseline estimate within 1-standard error.

[19]For each robustness check, the Probit model used to build $\hat{P}$ has been correspondingly adapted. Results for these Probit estimates (available upon request) are consistent with those reported. Table A3 in the Appendix also document the results obtained separately on each year. The estimated FC coefficient is positive in all years, showing that averaging variables over time only smooths the effect of individual-year specific shocks, while the main result is preserved.

[20]These countries are Austria, Belgium, Finland, France, Germany, Greece, Ireland, Luxembourg, Netherlands, Portugal, Spain.

2SLS-IV validation

We perform two further sets of validation exercises. First, we control for possible violations of the exclusion restriction. Second, we check whether our main result remains unchanged if we adopt different definitions of the FC dummy. Third, since a possible concern that our result is driven by a pure geographical effect may arise from the observed geographical distribution of our instrument, we estimate our baseline model on different geographical sub-regions. Results are presented in Table 5.

Checks on exclusion restrictions are as follows. In column 1, we estimate our baseline regression including additional provincial-level controls that might be correlated with the instrument and at the same time influence export price strategies of firms. As in Minetti and Zhu (2011), we add the (log of) value added and population at provincial level (provided by the Italian Statistical Office), and an index of infrastructural development of Italian provinces obtained from the research conducted by the Association of Italian Chambers of Commerce in collaboration with the "Guglielmo Tagliacarne" Institute. We find that the FC dummy coefficient is statistically equal to the baseline regression. Next, in column 2, we control for the financial level of development of the province at the beginning of the deregulation, by normalizing our instrument with the number of branches available in 1990. Also in this case the main result of a positive price premium for FC firms is preserved, and the magnitude is statistically comparable to the baseline regression.

Columns 3-4 present robustness to alternative definition of the FC group. We first estimate of the baseline model obtained with a weaker definition of the FC dummy, also including firms with rating equal to 7 in the FC group. Results (column 3) confirm again the positive price difference between constrained and unconstrained firms. The coefficient is larger (0.661) than in the baseline specification. This suggests that the group of rated-7 firms is indeed a borderline class, where a group of firms do behave as firms with more stringent financing problems: excluding them from the FC group underestimates the impact of FCs.[21] In column 4, we use a proxy of credit constraints commonly used in the financial constraints literature, that is firm leverage. We replicate the analysis by creating a FC dummy which takes value one for firms with leverage (assets-to-equity ratio) higher than the median, and zero otherwise. We still get consistent results.

Finally, we investigate the possible influence of spatial distribution of the instrument, by separate estimate of the baseline model for firms located in the North and in the Center-North of Italy. Results, reported in columns 5 and 6, respectively, show that there is some interplay with spatial factors, as indeed the FC coefficient is smaller than in the full sample baseline estimation. This suggests that FC firms are possibly more concentrated in the North. Nonetheless, the impact of FCs on prices remains positive and significant, lending support to our main conclusions.

4 Discussion

The strong message from the empirical analysis is that, controlling for firm characteristics, constrained firms sell at higher prices than unconstrained firms exporting within the same product-

[21] Note that the results are also consistent if we exclude the rated-7 firms from estimation of the baseline regression. The estimated γ is 0.396 in that case.

Table 5: FINANCIAL CONSTRAINTS AND (log) UNIT VALUE: IV-FE ROBUSTNESS CHECKS

	(1) Provincial Controls	(2) Initial development	(3) Weaker FC	(4) Leverage	(5) North	(6) North and Center
Second Stage results on (log) unit value						
Financially constrained firms dummy (FC)	0.296*** (0.034)	0.258*** (0.035)	0.661*** (0.020)	0.239*** (0.012)	0.102*** (0.032)	0.095*** (0.035)
log TFP	0.244*** (0.002)	0.254*** (0.002)	0.295*** (0.002)	0.262*** (0.002)	0.262*** (0.002)	0.249*** (0.002)
log number of employees	0.170*** (0.002)	0.168*** (0.002)	0.221*** (0.002)	0.176*** (0.002)	0.159*** (0.002)	0.162*** (0.002)
log age	0.015** (0.001	0.024*** (0.002)	0.074*** (0.002)	0.046*** (0.002)	0.025*** (0.001)	0.023*** (0.001)
log total assets	-0.127*** (0.002)	-0.122*** (0.002)	-0.203*** (0.003)	-0.124*** (0.001)	-0.124*** (0.002)	-0.114*** (0.002)
log gross operating margin	-0.038*** (0.001)	-0.042*** (0.002)	0.000 (0.002)	-0.033*** (0.001)	-0.045*** (0.001)	-0.046*** (0.002)
North dummy	0.197*** (0.004)	0.133*** (0.004)	0.099*** (0.004)	0.085*** (0.005)		
Center dummy	0.076*** (0.005)	0.130*** (0.004)	0.090*** (0.005)	0.078*** (0.005)		
log provincial value added	0.014*** (0.005)					
log provincial population	-0.063*** (0.006)					
Infrastructure Index	0.002*** (0.001)					
Panel B - First Stage on FC						
Instrument ($\hat{P}$)	0.990*** (0.007)	3.533*** (0.019)	0.724*** (0.007)	0.563*** (0.006)	0.561*** (0.006)	0.556*** (0.006)
Number of observations	2,454,168	2,422,274	1,904,266	2,343,791	2,454,168	2,454,168
Product-Country FE	Yes	Yes	Yes	Yes	Yes	
3 dgt sectoral dummies	Yes	Yes	Yes	Yes	Yes	

Notes: Second stage results, the dependent variable is (log) export unit value at the product-destination level. Column 1: additional provincial controls are added, i.e. (log of) value added, (log of) population and an index for the level of infrastructure. Column 2: we use as instrument the 1990-1998 average number of branches (per 1,000 inhabitants) divided by the number of branches (per 1,000 inhabitants) in 1990. Column 3: estimates with a less stringent definition of the FC dummy, including firms rated as 7. Column 4: the FC dummy is 1 for firms with leverage above the median, and 0 otherwise. Column 5: baseline model is estimated only on firms located in the North of Italy. Column 6: baseline regression only on firms located in the North and Center of Italy. All the regressions include a constant, HS6 product-country pair fixed effects and 3-digit industry dummies.

Robust standard errors in parenthesis, clustered at product-country level. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

destination market.

There are different mechanisms which can explain this finding. Industrial organization models of competition in markets with demand rigidities or imperfect consumer adjustment represent a first natural candidate. The logic in these models is that increasing prices is just a way, for firms unable to access external finance, to increase revenues in the attempt to keep financing investment and to meet current liabilities. However, some form of friction in the product market is needed to allow constrained firms to raise their prices without loosing too much demand, at least for a short period of time. Factors like brand loyalty or, more in general, the existence of substantial switching costs for buyers in moving from one seller to another are typical examples of such frictions (see the survey in Klemperer, 1995). Alternatively, one can think about models like Gottfries (2002) or Lundin et al. (2009), based on the classic customer market model (Phelps and Winter, 1991), which directly postulate a time lag in consumers' reaction to price changes. Instead, Gagnon (1989) assumes the existence of large adjustment costs for changing quantities, so that acting on prices is the only viable strategy for constrained firms in the short run.

To illustrate the key mechanism relating prices and capital market imperfections, consider a simple duopoly where two firms compete over two periods producing the same good with equal marginal costs (or productivity). Assume that there exists a certain degree of product differentiation, and that consumers have to pay a cost if they are willing to switch seller from one period to another.[22] Such switching costs create an incentive for firms to price below the standard marginal cost pricing rule. Indeed, in period 1, firms find it optimal to lock-in an higher number of consumers via a price reduction, anticipating that they will be able to charge a monopoly price in the second period. To see the impact of financing problems in this setting, assume that a firm cannot entirely cover operations or investment via internal resources. The same firm has, however, access to external finance through a standard incentive compatible contract under the assumption that the true cash flow is not observed by the creditor institution, but observed and maybe be diverted to perks or pet projects by managers. The optimal contract foresees full-repayment at end of period 1, and it imposes a threat of full liquidation at end of period 1 if the repayment is not met.[23] This creates a weaker incentive to keep prices below standard marginal costs: the gains from locked-in consumers are lower in expectations, as there is now a positive probability that the firm will not meet the contract. Given strategic interaction, both constrained and unconstrained firms set a price higher than the price without capital markets imperfections, but the upward distortion is stronger for constrained firms.[24]

[22] One can model product differentiation by means of a unit transportation cost, assuming as standard that this represents an opportunity cost payed by consumers if a product in the market does not match her/his ideal variety.

[23] See Chevalier and Scharfstein (1996). Notice that liquidation is an inefficient outcome since firm assets are worth less if managed by the investors.

[24] In principle, one can also imagine models where constrained firms choose aggressive pricing (i.e. sell at lower price than competitors), seeking to sustain profits by expanding market shares. Models of pricing under financial distress (see for instance Dasgupta and Titman, 1998; Pichler et al., 2008) show that the incentive to raise rather than reducing prices might depend on the maturity structure of debt: the higher the burden of short term repayment, the higher the discount of future revenues as compared to current revenues, and the higher the probability that firms set higher prices and renounce to market shares in the short run. However, our findings suggest that this is not the case in practice, possibly because building the capacity to accommodate the increase in demand requires extra finance that constrained firms are typically not able to access. This is in line also with the standard result in the literature that constrained exporters export less than unconstrained exporters.

Within the international trade literature there are few attempts (Chaney, 2013; Muuls, 2008; Manova et al., 2011) to incorporate financing constraints in the monopolistic competition model with heterogeneous firms proposed in Melitz (2003). The recent work by Fan et al. (2012) provides a detailed focus on the effects of FCs on export prices, at the same time sharing the basic features of other models. The model setting assumes consumers with a standard quality-adjusted CES utility function over different varieties of the same good, which entails that, in addition to the standard price-quantity relation, higher quality and more advertised products generate larger demands. On the production side, firms are heterogeneous in productivity and each produce a specific variety of the good with a single input (i.e. labor) technology. Firms face an iceberg trade cost and, following a common approach, the marginal cost of production is an increasing function of quality: higher quality implies higher marginal costs, induced by the hiring of higher quality workers or by investing in R&D. Moreover, each firm faces two different types of fixed costs: a production fixed cost capturing the fixed investment in quality and a marketing fixed cost, modeled as a function of the advertisement intensity, capturing all the costs of penetrating into foreign markets. The role of financial constraints is modeled similarly to Manova (2013), so that firms need external capital to finance a fraction of all types of costs (fixed and variables).[25] However, due to financial constraints, firms can borrow only up to a fraction of their cash flow. In this set up two mechanisms connect prices to financing problems. First, controlling for productivity, constrained firms sell at higher prices than unconstrained firms to sustain cash flow and lessen the constraints. This mechanism is called "price distortion" effect. Here advertisement plays a role similar to consumers' switching costs or any other source of demand rigidity creating the scope for frictional response of consumers to higher prices. However, there is also a second, "quality adjustment", effect which pushes prices of constrained firms down. Again controlling for productivity, financially constrained firms sell lower quality goods at lower prices since they cannot afford the higher cost of quality. Fan et al. (2012) show that when product quality is endogenously determined the latter effect dominates on the former causing financially constrained firms to charge lower prices than unconstrained ones. On the contrary, when quality is exogenous constrained firms export at higher prices.

Based on these theoretical considerations, the positive effect of financial constraints on export prices that we observe in the data is compatible with two alternative explanations. First, that prices raise because constrained firms revise upward their mark-up over given costs (i.e. there exists only a price distortion effect). Or, second, constrained firms set higher prices because the cost reduction induced by the choice of producing lower quality goods is crowded out by an increase in the corresponding mark-up (i.e. quality adjustment and a price distortion effects coexist, but the latter dominates the former). In addition to these two channels, and regardless the type of model considered, there might be a third simple explanation based on costs. Indeed, if it is the case that constrained firms bear higher unit costs, irrespective of the quality level of their products, then they might charge higher export prices than unconstrained firms just because of a standard marginal cost pricing rule.

[25] Predictions do not change if external finance is needed to only cover fixed costs.

Table 6: MARGINAL COSTS, MARK-UP, QUALITY

	(1) Marginal Costs	(2) Low HD	(3) High HD	(4) Low VD	(5) High VD
Panel A - Second Stage on (log) unit value					
Financially constrained firms dummy (FC)	0.331*** (0.034)	0.019*** (0.044)	0.671*** (0.056)	0.564*** (0.050)	0.051 (0.046)
log TFP	0.219*** (0.002)	0.261*** (0.004)	0.192*** (0.003)	0.219*** (0.004)	0.243*** (0.004)
log number of employees	0.164*** (0.002)	0.140*** (0.003)	0.205*** (0.004)	0.212*** (0.003)	0.106*** (0.003)
log age	0.018*** (0.001)	0.033*** (0.002)	0.016*** (0.003)	0.011*** (0.002)	0.040*** (0.002)
log total assets	-0.121*** (0.002)	-0.094*** (0.003)	-0.153*** (0.004)	-0.162*** (0.003)	-0.072*** (0.004)
log gross operating margin	-0.036*** (0.001)	-0.052 (0.002)	-0.021*** (0.003)	-0.030*** (0.002)	-0.044*** (0.002)
log unit wage	0.063*** (0.002)	0.020*** (0.003)	0.074*** (0.003)	0.052*** (0.003)	0.048*** (0.004)
log interest rate	0.033*** (0.001)	0.020*** (0.002)	0.052*** (0.002)	0.039*** (0.002)	0.032*** (0.002)
North dummy	0.131*** (0.004)	0.115*** (0.007)	0.135*** (0.006)	0.128*** (0.005)	0.134*** (0.008)
Center dummy	0.125*** (0.005)	0.128*** (0.008)	0.133*** (0.007)	0.153*** (0.006)	0.110*** (0.009)
Panel B - First Stage on FC					
Instrument ($\hat{P}$)	0.668*** (0.008)	0.846*** (0.012)	0.670*** (0.013)	0.655*** (0.012)	0.877*** (0.012)
log TFP	-0.022*** (0.001)	-0.005*** (0.001)	-0.033*** (0.001)	-0.023*** (0.001)	-0.014*** (0.001)
log number of employees	-0.010*** (0.0003)	-0.007*** (0.0004)	-0.014*** (0.0004)	-0.009*** (0.0004)	-0.012*** (0.0004)
log age	-0.007*** (0.0002)	-0.002*** (0.0003)	-0.014*** (0.001)	-0.013*** (0.0003)	-0.002*** (0.0004)
log total assets	0.020*** (0.0003)	0.014*** (0.0004)	0.028*** (0.001)	0.022*** (0.0005)	0.018*** (0.0005)
log gross operating margin	-0.004*** (0.0003)	0.000 (0.0007)	-0.007*** (0.001)	-0.008*** (0.001)	0.002** (0.0005)
log unit wage	0.003*** (0.0004)	-0.007*** (0.001)	0.011*** (0.001)	0.006*** (0.001)	-0.003*** (0.0007)
log interest rate	0.003*** (0.0002)	0.001*** (0.0002)	0.004*** (0.0002)	0.001*** (0.001)	0.004*** (0.002)
North dummy	0.006 (0.001)	-0.000 (0.001)	0.013*** (0.001)	0.008*** (0.001)	0.006*** (0.001)
Center dummy	0.027*** (0.001)	0.023*** (0.001)	0.040*** (0.001)	0.031*** (0.001)	0.033*** (0.001)
Number of observations	2,139,625	924,740	834,130	933,614	825,256
Product-Country FE	Yes	Yes	Yes	Yes	Yes
3 dgt sectoral dummies	Yes	Yes	Yes	Yes	Yes

Notes: Panel A - second stage results, the dependent variable is (log) export unit value at the product-destination level. Panel B - first stage results, the dependent variable is the FC dummy. In column 1 we add to the baseline model the (log of) unit wage and the (log of) firm interest rate (interest expenses over stock of financial debt). In columns 2 and 3 regressions are estimated separately for HS6 products with below or above median values of horizontal differentiation (based on the dissimilarity component of Gallop-Monahan index), respectively. In columns 4 and 5 regressions are estimated separately for HS6 products with below or above median values of vertical differentiation (based on R&D plus advertisement share of sales, from in Kugler and Verhoogen, 2012), respectively.

Robust standard errors in parenthesis, clustered at product-country level. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Marginal Costs, Mark-up, Quality

In line with the theoretical framework discussed above we conclude the paper by performing further empirical investigations to disentangle the relative importance of the three different interpretations, at the same time allowing for additional qualifications of our main finding.

Accounting for the marginal cost effect requires to enrich our baseline regression model to include firm-level proxies of unit costs among the regressors, and test whether the coefficient on the FC dummy remains positive and statistically significant. Specifically, we add the (log of) unit wage, obtained as the ratio between the total labour expenses and the number of employees, and the (log of) "firm interest rate", proxied by the ratio between yearly interest expenses and the stock of financial debts. 2SLS-IV results are reported in column 1 of Table 6. We find that our FC dummy still has a positive and significant coefficient. As expected, the unit wage and the interest rate are positively associated with firms' export prices, while results on the other controls are in line with the baseline regression. Constrained firms are confirmed to charge higher prices, therefore, even after controlling for the possible marginal cost mechanism.

Next, we try to improve our understanding of the relative importance of the mark-up or price-distortion effect vs. the quality adjustment mechanisms, by comparing patterns of estimates across horizontally and vertically differentiated products. The rationale behind the exercises is the following. First, irrespective of the existence of quality adjustments, when the varieties of a good provided to the market are highly horizontally differentiated, it is easier for the competing firms to act on prices without drastically affecting their market shares. Accordingly, a stronger mark-up or price-distortion effect, and thus an higher coefficient on the FC dummy, should be observed within export transactions involving less substitutable products. Second, the quality adjustment mechanism, leading constrained firms to export lower quality at lower price, should be stronger in products with high vertical differentiation, where there is greater scope to act on quality differences, than in products where there is narrower or no scope at all for quality up/down-grading. Thus, if adjustments related to quality choices do matter, then the coefficient on the FC dummy should be smaller in products where vertical differentiation is more important.

To measure the degree of horizontal differentiation of exported products we exploit the classification constructed by Kugler and Verhoogen (2012) using Colombian firm-level data. It is based on dissimilarity component of the full Gallop-Monahn index, which measures dissimilarity of input mixes across firms within an industry.[26] Industries characterised by an high degree of dissimilarity are those producing products which are less substitutable and thus more horizontally differentiated.[27]. To measure vertical differentiation of the export product categories we again follow Kugler and Verhoogen (2012) and employ their classification based on the ratio of advertising plus R&D expenditures to total sales in U.S. industries.[28] The logic here is that firms invest in R&D and advertising in sectors

[26]This is the distance between the expenditure share on each input i in total expenditure of a firm active in a certain industry and the average expenditure share on input i by all firms in the same industry. The sum of these distances weighted by the revenue share of the firm in total industry revenues yields the industry-level measure.

[27]We convert the original ISIC (Rev.2) 4-digit level classification of sectors into HS6 product level exploiting the concordance tables from http://www.macalester.edu/research/economics/page/haveman/Trade.Resources/tradeconcordances.html

[28]The original data are from the U.S. Federal Trade Commission 1975 Line of Business Survey. Kugler and Verhoogen (2012) convert FTC 4-digit industry classification into ISIC (Rev. 2) 4-digit classification using verbal industry

wherein it is possible to affect quality and thus there is scope for quality differentiation.[29]

We re-estimate equation (1) with 2SLS-IV, separately on different sub-samples of export transactions involving products with different degrees of horizontal or vertical differentiation. Notice that we still include the proxies of unit costs, so that the results hold for given marginal costs effect.

Table 6 presents the estimates. Columns 2 and 3 explore the impact of horizontal differentiation. The two groups of High and Low differentiated products are distinguished as the HS6 categories with dissimilarity index above and below the median computed across all HS6 product categories. The coefficient estimates for the FC dummy are still positive in both groups, but the estimated impact is higher across firms operating in highly differentiated products. This confirms the expectation of a stronger positive price premium for constrained firms active in more differentiated sectors. The point estimate is 0.671, about twice as larger than the unsplitted estimates in column 1. This magnitude suggests that horizontal differentiation, and the related scope for price-distortion or mark-up effects related to locking-in customers, are indeed crucial drivers of pricing strategies.

In columns 4 and 5 of Table 6 we instead compare 2SLS-IV results across two groups of products with High or Low vertical differentiation. These are defined as HS6 product categories where the vertical differentiation index is above or below the median value. We observe that the FC dummy coefficient is smaller and not significant in the sub-sample relative to more vertically differentiated products. This evidence is in agreement with the existence of a quality-adjustment mechanism. Indeed it tells that, when there is more room for quality adjustment, the positive mark-up or price-distortion effect is counteracted and actually offset by a reduction of the price compatible with constrained firms choosing to export lower quality products at lower prices.[30]

5 Conclusion

The present paper provides a comprehensive analysis of the role that financial constraints play in shaping firms' export prices. We use detailed firm-product-country data on the international activities of a sample of firms covering the vast majority of Italian exporters. We are able to relate export prices directly to firm-level credit status, by exploiting a firm-level information on access to credit based on credit ratings issued by an independent institution. Moreover, we overcome an important limitation of previous studies, by adopting an instrumental variable strategy which allows a proper identification of the effect of credit constraints. Thus, our article is the first to establish a causal link from the financial status of exporters to their pricing strategies.

Our key findings tell that firms facing tighter credit conditions charge higher prices, even after controlling for key firm attributes, including size, productivity, financial variables, and product-

descriptions. We convert from ISIC 4 digit level to HS6 product level using the appropriate concordance tables.

[29] As an additional robustness check, we also employ the Rauch (1999) measure, based on whether a good is traded on a commodity exchange or it has quoted price in industry trade publications. This measures "overall" differentiation (i.e. both horizontal and vertical). In fact, as argued by Kugler and Verhoogen (2012), although the trade literature has extensively used the Rauch index as a measure of horizontal differentiation, it is indeed unclear which dimension it proxies for. Results are available upon request.

[30] We performed additional exercises investigating the role of quality adjustment, either including a proxy of input quality in the baseline regression, or estimating the model separately for transactions with developed vs. developing destination countries. The positive coefficient on the FC dummy is preserved, see Table A4 in the Appendix for details.

destination fixed effects. This positive price premium of financial constraints is shown to be robust to a series of sensitivity analyses. It holds true even if we remove firms with only one employee, drop product and destination countries representing negligible shares of a firm's exports, exclude multinational firms, or exclude those transactions operated with a currency different from the Euro. Also, the main message remains unchanged if we experiment with different definitions of the group of constrained firms, or play with the degree of local variation of the data which is key for identification given the provincial-level variability that we exploit in building our instruments. Finally, the main result still come up if we try to assess its relationship with mark-up strategies and quality, by adding measures of unit costs and by looking at horizontal vs. vertical differentiation.

Our study contributes to the literature documenting the systematic variation in export prices across firms, products and trade partners. The results point out that financial constraints, in addition to efficiency or quality, should be taken into account to explain export price differences across firms. In this respect, our main finding posit a challenge for the theory since it points at the co-existence of two mechanisms influencing export prices under financial constraints. A price distortion or mark-up effect that tends to push the price charged by constrained firms up, and an opposing quality adjustment mechanisms through which constrained firms can set lower price. However, the former effect dominates the latter. Existing models predict such a positive price premium for constrained firms only when quality is exogenous, an assumption that it is hardly met in reality and already disproved by empirical studies. As a result, non of the existing theoretical explanations appears able to reconcile the ensemble of our empirical findings.

Appendix

The structure of the Italian financial system

The structure of the Italian financial system presents some peculiarities with respect to other major countries. Figures A1-A3 in this Appendix show that in Italian firms' external financing mainly occurs through banks, partly due to the underdevelopment of bond and stocks markets.

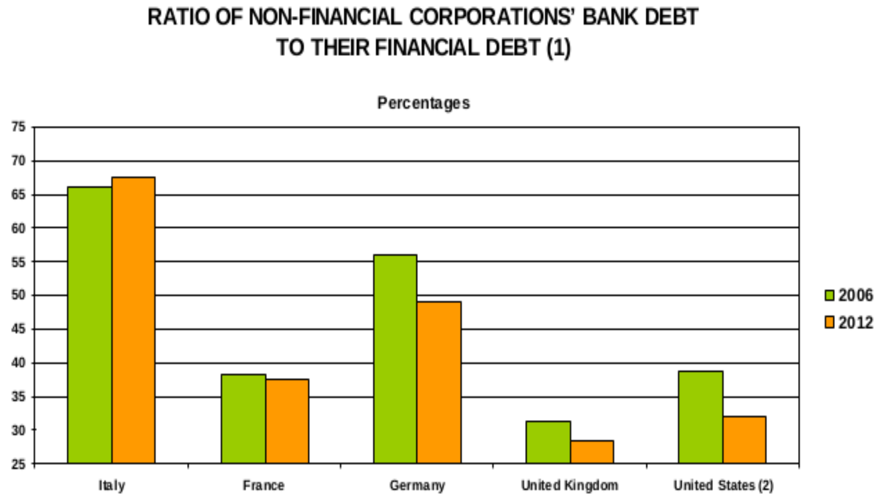

Figure A1. Source: Panetta (2013). Based on data taken from Bank of Italy for Italy; Eurostat and ECB for the euro-area countries; Bank of England for the United Kingdom; Federal Reserve System for the United States. (1) Bank debt comprises only the loans disbursed by the banks resident in each country; (2) 2011 data.

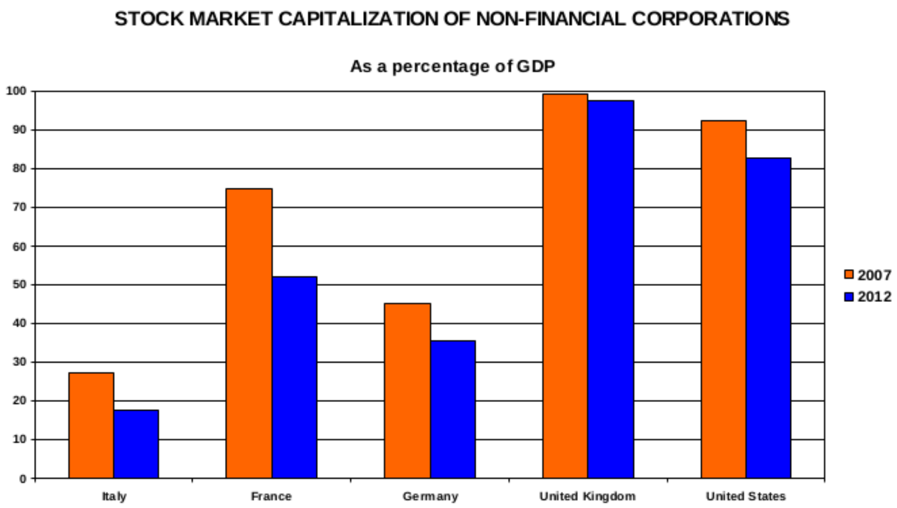

Figure A2. Source: Panetta (2013). Based on Datastream data.

Custom data

In compliance with the common framework defined by the European Union (EU), there are different requirements in order for a cross-border transaction to be recorded, depending on whether the importing partner is an EU or NON-EU country, and on the value of the transaction.

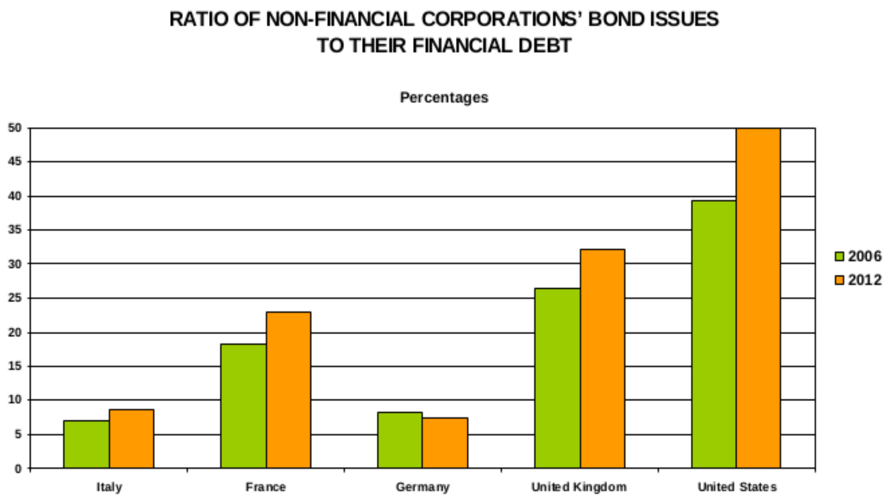

Figure A3. Source: Panetta (2013). Based on data taken from Bank of Italy for Italy; Eurostat and ECB for the euro-area countries; Bank of England for the United Kingdom; Federal Reserve System for the United States.

As far as outside EU transactions are concerned, there is a good deal of homogeneity among member states as well as over time. In the Italian system the information is derived from the Single Administrative Document (SAD) which is compiled by operators for each individual transaction. From the introduction of the Euro, Italy has set a threshold at 620 euro (or 1000 Kg) for a transaction to be recorded. For all of these recorded extra-EU transactions, the COE data report complete about product category, destination, quantity and value.

Transactions within the EU are collected according to a different system (Intrastat). There the thresholds on the value of transactions qualifying for complete record are less homogeneous across EU member states, with direct consequences on the type of information reported in the data. In 2003 (the last year covered in the analysis), there are two cut-offs. If a firm has more than 200,000 euro of exports (based on previous year report), then she must fill the Intrastat document monthly. This implies that complete information about product types is also available. Instead, if previous year export value falls in between 40,000 and 200,000 euro, the quarterly Intrastat file has to be filled, implying that only the amount of export is recorded, while information on the product is not. Firms with previous year exports below 40,000 euro are not required to report any information on trade flows. According to ISTAT, although only one-third of the operators submitted monthly declarations, these firms cover about 98% of trade flows (http://www.coeweb.istat.it/default.htm). Thus, firms which do not appear in COE are either marginal exporters or do not export at all.

Representativeness

Table A1 shows that the representativeness of the dataset is quite satisfactory. We report here 2003 data, but figures are comparable in the other years. As mentioned in the main text, although the dataset includes only about 20% of manufacturing in terms of number of firms, we cover about 60% of manufacturing firms that do export, and about 84% of the total value of manufacturing exports. We see here that these number are also fairly stable across different industrial sectors.

This picture is explained by the well known abundance of micro and small firms in Italian manufacturing, together with the observation that the legal status of limited firm tend to be more spread across medium-bigger firms, and medium-big firms are expected to account for the great bulk of overall export activities in the country, in line with a well established result in the literature. In agreement with this, Table A2 shows that, again for 2003 but valid across other sample years, the firms in our sample are on average slightly bigger and more productive (in terms of labour productivity here) than the population of manufacturing firms. At the same time, however, we do not observe big differences when we focus on exporting firms: the average size, labour productivity, export values, number of exported products and number of destinations served do not differ significantly between our sample and the population.

Table A1. COVERAGE OF THE DATASET, MANUFACTURING: NUMBER OF FIRMS; NUMBER OF EXPORTERS, EXPORT VALUE (2003)

	ALL FIRMS			EXPORTERS			EXPORT VALUE		
Sector	ASIA-COE (Number)	Our dataset (Number)	Coverage %	ASIA-COE (Number)	Our dataset (Number)	Coverage %	ASIA-COE (billion)	Our dataset (billion)	Coverage %
15	71345	8882	12.45	4927	2872	58.36	12.1	9.4	77.77
17	27762	6408	23.08	5681	3445	60.69	12.5	10.8	86.70
18	41615	6134	14.74	5035	2654	52.73	9.7	8.1	83.56
19	21985	4495	20.45	5688	2644	46.48	10.8	8.8	81.62
20	46584	3550	7.62	2458	978	39.79	1.5	1.3	83.88
21	4566	1951	42.73	1328	884	66.57	4.0	3.8	95.28
22	27344	7801	28.53	2164	1237	57.26	1.7	1.6	91.25
23	443	333	75.17	84	72	86.90	3.8	3.7	99.25
24	6127	3529	57.60	2595	1984	76.61	22.6	16.3	71.80
25	13084	5575	42.61	4422	2968	67.18	10.4	8.9	85.72
26	27230	6218	22.84	4522	2176	48.12	7.2	6.2	86.18
27	3814	1893	49.63	1335	1016	76.10	9.9	8.7	88.21
28	99519	19551	19.65	10280	5754	56.17	12.6	11.2	89.26
29	42391	14710	34.70	12128	8177	67.55	43.3	38.0	87.61
30	1976	822	41.60	262	185	70.61	1.5	1.3	91.19
31	18316	5315	29.02	3214	2128	66.30	8.1	6.6	82.12
32	8671	1665	19.20	911	608	66.85	5.2	3.7	71.02
33	22399	3073	13.72	1921	1355	70.68	4.6	3.9	85.18
34	1962	1122	57.19	918	687	74.84	17.8	15.3	85.86
35	4684	1541	32.90	819	475	60.81	6.7	4.9	73.84
36	50018	7873	15.74	8664	4193	48.42	12.1	10.4	85.96
Total	541835	112441	20.75	79356	46492	58.69	218.1	183.0	83.93

Notes: The Table reports, for 2003, the number of firms, the number of exporters and the export value by sector for the entire population of Italian manufacturing firms (ASIA-COE dataset) and the limited liabilities firms (our dataset).

Table A2. DESCRIPTIVE STATISTICS, 2003

	ASIA-COE			Our Dataset		
	Mean	Sd	Observations	Mean	Sd	Observations
	Manufacturing firms					
log number of employees	1.12	1.14	541836	2.13	1.38	112441
log Total Sales/num.employees	3.78	1.12	518839	4.65	1.09	110160
	Manufacturing Exporters					
log number of employees	2.43	1.35	79352	2.85	1.32	46574
log Total Sales/num.employees	11.74	0.94	77068	11.99	0.82	46073
log export	4.71	2.74	79352	5.52	2.67	46574
log number of destinations	8.77	12.92	79352	11.66	14.74	46574
log number of products	8.04	14.7	79352	10.36	17.15	46574

Notes: The Table reports, for 2003, some descriptive statistics for the entire population of Italian manufacturing firms (ASIA-COE dataset) and for limited liabilities firms (our dataset).

Probit Goodness of fit

The 2SLS-IV estimator employed in the paper follows Procedure 21.1 in Wooldridge (2010). To recall, the procedure entails a two stage estimator for the dummy endogenous variable model in equation 1. First, estimate by maximum likelihood the binary response model $P(\text{FC}_f = 1|\mathbf{X},\text{Z})$ where $\mathbf{X}$ are firm level controls and Z the instrument, and obtain the associated fitted probabilities $\hat{P}$; Second, estimate equation 1 by 2SLS-IV using the fitted probabilities $\hat{P}$ as instrument.

As explained in the main text, to build $\hat{P}$ we estimate by maximum likelihood the following probit

$$\text{P}(\text{FC}_f\text{=1} \mid \mathbf{X},\text{Z}) = \Phi\left(\delta_1 \text{Z} + \mathbf{X}'_f \boldsymbol{\beta} + \epsilon_f\right) \quad . \tag{3}$$

where the probability of being financially constrained is regressed on Z, the average number of new branches opened annually between 1990-98, and on $\mathbf{X}$, the set of firm level controls.

Although this IV estimator enjoys several good properties, the standard weak instrument diagnostics fail in this context. We provide here a series of alternative assessment of the goodness of fit of the probit, supporting the validity of the instrument.

Figure A4 reports (upper panel) separation plot (Greenhill et al., 2011), a simple visual technique for assessing the predictive power of models with binary outcomes. Its main advantage is that it does not depend on the often arbitrary choice of a probability threshold to distinguish between predicted positive and negative events. The separation plot reports on the x-axis the fitted probabilities of observing a positive event $\hat{\text{P}}$(FC=1) and it associates to each value of $\hat{P}$ the actual value of the dummy FC. The starker the separation between the colored and uncolored areas the better the goodness of fit of the model. For comparison, the bottom panel of Figure A4 reports the separation plot for a model with only age as control, showing the improvement in the prediction ability one can obtain by adding relevant controls.

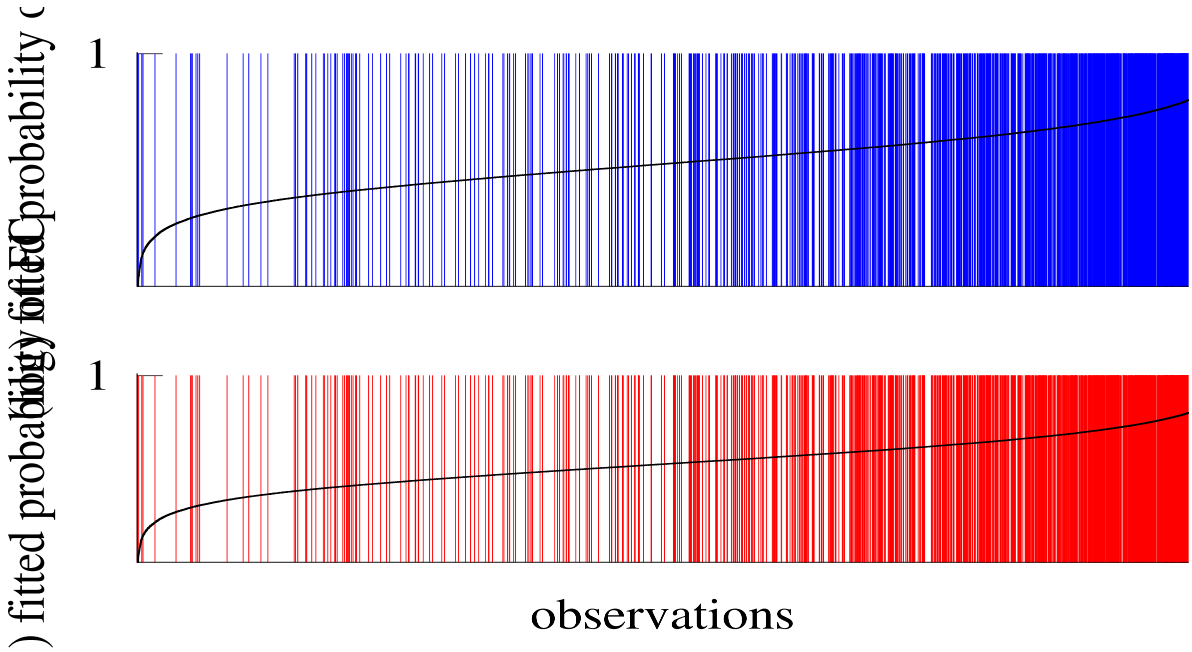

Figure A4: PROBIT SEPARATION PLOT

Panel B of Table 3 includes three other goodness of fit statistics. Together with a standard Adjusted-R^2, we report the Brier and AUC scores. The former is an index of prediction error for binary event calculated as $BS = (1/N)\sum(\hat{FC} - FC)$ where $\hat{FC}$ is status predicted by the model and FC the actual status. The latter measures the area under the ROC curve (Receiver Operating Characteristics), yet another estimate of the predictive power of the model. Indeed, the ROC curve represents the relation between the false positive rate (FPR) of the model FP/(FP+TN) and its true positive rate (TPR) that is TP/(TP+FN). Ideally, one would like to have a model with a TPR equal to 1 and a FPR equal to 0. In practice, the closer the ROC curve to the border of the unit-length square, the higher the AUC score and the higher the ability of the model to forecast the event under scrutiny. The ROC curve for model 3 is reported in Figure A5.

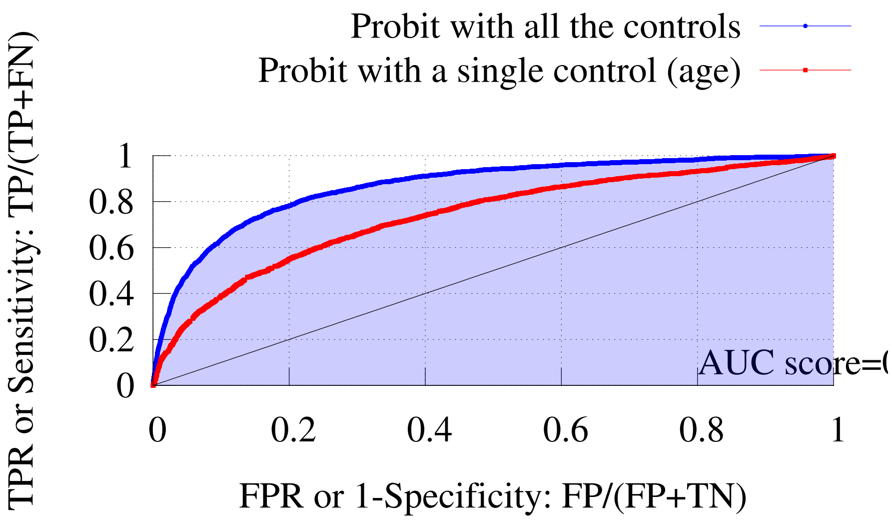

Figure A5: ROC PLOT

Additional robustness checks

We provide here additional robustness checks that complement the results presented in the main text.

Table A3 reports results of estimates of the baseline regression model

$$\ln \text{EUV}_{fpc,t} = \gamma \text{FC}_{f,t-1} + \mathbf{X}'_{f,t-1}\boldsymbol{\beta} + \mu_{pc} + \epsilon_{fpc,t} \quad , \tag{4}$$

performed separately on each year. The FC dummy coefficient is positive and significant in all years. This shows that averaging the data over time has the only effect to smooth yearly-specific shocks, while the main result is preserved.

In Table A4 we present further analysis of the role of quality adjustment, complementing Section 4. First, we include a control for firm input quality. Following Manova and Zhang (2012a), we measure firm input quality via a weighted average of the unit values of all the import transactions in intermediate inputs performed by a firm. The FC coefficient is still positive and significant. Second, we split the sample between export transactions to developed vs. to developing countries. The coefficient on FC is smaller in developed countries, where consumers are more likely to consume high quality products and therefore there is higher scope for quality differentiation.

Table A3 - FINANCIAL CONSTRAINTS AND (log) UNIT VALUE: IV-FE BY YEAR

	(1) 2001	(2) 2002	(3) 2003
Second Stage on (log) unit value			
Financially constrained firms dummy (FC)	0.627*** (0.35)	0.783*** (0.078)	0.486*** (0.032)
log TFP	0.279*** (0.003)	0.299*** (0.003)	0.296*** (0.003)
log number of employees	0.178*** (0.003)	0.184*** (0.002)	0.180*** (0.002)
log age	0.038*** (0.002)	0.035*** (0.002)	0.040*** (0.002)
log total assets	-0.151*** (0.002)	-0.150*** (0.002)	-0.145*** (0.002)
log gross operating margin	-0.023*** (0.001)	-0.032*** (0.001)	-0.041*** (0.00)
North dummy	0.101*** (0.006)	0.106*** (0.006)	0.116*** (0.005)
Center dummy	0.100*** (0.006)	0.103*** (0.006)	0.114*** (0.006)
Number of observations	1,091,249	1,196,947	1,392,799
Product-Country FE	Yes	Yes	Yes
3 dgt sectoral dummies	Yes	Yes	Yes

Notes: Second stage results, the dependent variable is (log) export unit value at the product-destination level. Regression in column 1 estimated using 2001, in column 2 using 2002, and in column 3 using 2003. All the regressions include a constant, HS6 product-country pair fixed effects and 3-digit industry dummies.

Robust standard errors in parenthesis, clustered at product-country level. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Table A4: ROLE OF QUALITY, IV-FE ROBUSTNESS CHECKS

	(1) Input Quality	(2) Developed	(3) Developing
Panel A - Second Stage on (log) unit value			
Financially constrained firms dummy (FC)	0.212*** (0.037)	0.313*** (0.038)	0.570*** (0.093)
log TFP	0.191*** (0.002)	0.216*** (0.003)	0.258*** (0.008)
log number of employees	0.157*** (0.002)	0.164*** (0.002)	0.170*** (0.005)
log age	0.014*** (0.001)	0.014*** (0.002)	0.050*** (0.004)
log total assets	-0.113*** (0.002)	-0.123*** (0.002)	-0.123*** (0.006)
log gross operating margin	-0.036*** (0.001)	-0.036*** (0.002)	-0.034*** (0.004)
log unit wage	0.066*** (0.002)	0.063*** (0.002)	0.074*** (0.007)
log interest rate	0.042*** (0.001)	0.032*** (0.001)	0.032*** (0.004)
Import Average Price	0.444*** (0.002)	0.133*** (0.004)	0.128*** (0.011)
North dummy	0.105*** (0.004)	0.133*** (0.004)	0.128*** (0.011)
Center dummy	0.105*** (0.005)	0.129*** (0.005)	0.087*** (0.013)
Panel B - First Stage on FC			
Instrument ($\hat{P}$)	0.646*** (0.008)	0.649*** (0.008)	0.791*** (0.022)
Number of observations	2,027,245	1,767,762	278,353
Product-Country FE	Yes	Yes	Yes
3 dgt sectoral dummies	Yes	Yes	Yes

Notes: Panel A - second stage results, the dependent variable is (log) export unit value at the product-destination level. Panel B - first stage results on the instrument $\hat{P}$, dependent variable is Pr(FC=1). Column 1: we add as regressor the firm level average unit value of imports of intermediate inputs across products and destinations. Columns 2 and 3: separate regressions for the groups of developed and non-developed destination countries, respectively. All the regressions include a constant, HS6 product-country pair fixed effects and 3-digit industry dummies.

Robust standard errors in parenthesis, clustered at product-country level. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

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