Financial constraints and firm exports: accounting for heterogeneity, self-selection and endogeneity

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Abstract

The paper examines the causal effect of financial constraints on firms' exports. We exploit a firm-level proxy of constraints based on credit ratings and available for a large panel of Italian exporting and non exporting firms. Our estimation strategy allows to cure for self-selection into exports and endogeneity of financial constraints. At the same time, we can control for unobserved firm fixed effects both in the selection and in the export equation, thus identifying the effect on exports of within firm changes in financial constraints status. We find that financial constraints produce a sizable reduction in the value of a firm's foreign sales.

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

Keywords: financial constraints, exports, self-selection, endogeneity

1 Introduction

A rapidly growing body of research in the trade literature examines the role of external finance in shaping the international activities of firms. Selling to foreign markets indeed involves specific fixed and variable costs, which are additional to those required to operate in the domestic market and may thus create a specific need to resorting to external credit.[1] Theoretical studies (see Chaney, 2005; Muuls, 2008; Manova, 2011; Feenstra et al., 2011) incorporate this idea within the standard Melitz (2003) model of international trade with heterogeneous firms. In spite of differences in modeling financial constraints, all the theoretical frameworks share the common prediction that financing problems reinforce self-selection into export markets driven by productivity. Indeed, the productivity level required to enter and operate in international markets under financial constraints is higher than in the absence of constraints because firms must also cover the costs of external finance. If external credit is needed to only meet sunk and fixed costs of exports, then financial constraints are predicted to only affect the probability to become an exporter (i.e. the extensive margin), with constrained firms less likely to enter foreign markets. If, conversely, external funds are needed to cover both fixed and variable export costs, then financial constraints also affect the overall value of foreign sales (i.e. the intensive margin): *ceteris paribus*, constrained firms that enter foreign markets export less than unconstrained exporters.

The existing empirical literature is usually interpreted as supporting these predictions, although there are exceptions casting doubts on whether the available evidence can be considered as conclusive. Concerning entry into export markets, financial constraints are found to reduce the probability to become exporter in Muuls (2008) for Belgium, in Bellone et al. (2010) for France, in Wagner (2012) for Germany, in Berman and Hericourt (2010) for a sample of nine developing and emerging economies, 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. On the contrary, constraints do not significantly affect entry into exports according to the evidence in Greenaway et al. (2007) for the UK, Stiebale (2011) for France and Arndt et al. (2012) for Germany. Most of the studies, with the notable exception of Muuls (2008), Berman and Hericourt (2010) and Arndt et al. (2012), also find that financial constraints affect the intensive margin of exports by reducing the value of a firm's exports. Further, Damijan et al. (2010) find that improving the access to external finance helps Slovenian firms to expand exports and even more so for small firms.

In this paper, exploiting a large representative panel of Italian firms, we propose a further look at the empirical identification of the effect of financial constraints on firms' foreign sales. We address what we perceive as important limitations in previous studies, providing two distinct contributions.

First, we jointly consider a fixed effect type of control for unobserved firm heterogeneity together with two well-known sources of potential bias, namely self-selection into exports and potential endogeneity of financial constraints in the determination of firms' exports. Among previous empirical papers, Minetti and Zhu (2011) make the only attempt to address self-selection and endogeneity at the same time, notably on a sample of Italian firms smaller than the sample available to us. They employ

[1] Beck (2002) and Svaleryd and Vlachos (2005), among others, provide evidence of a negative effect of financial development on aggregate exports of countries and sectors.

a modified Heckman-type procedure to deal with selection, and exploit exogenous variation in the geographical distribution of local supply of banking services to instrument their proxy of credit constraints. Though, their cross-sectional data do not allow to control for unobserved heterogeneity. Our main step forward with respect to the literature is precisely along this direction. By applying the estimator developed in Semykina and Wooldridge (2010), we are able to fully exploit the panel dimension of our data to control for unobserved heterogeneity in both the selection and the main equation, and, at the same time, to allow for an instrumental variable treatment of potential endogeneity of access to credit. Thus, we identify the causal effect on exports of within firms changes in financial conditions, while previous studies only capture differences across constrained vs. unconstrained firms. A clear identification of this effect is of crucial importance as it contributes to the debate concerning proper policies to support and foster international expansion of manufacturing firms.

Second, a key issue concerns the intrinsic difficulty in measuring financial constraints. Theoretically the goal is clear: finding an empirical indicator identifying the point where the credit supply curve faced by a firm becomes inelastic. However, the empirical measures adopted in the literature, inspired by a long standing debate outside the trade literature, are many and different. This heterogeneity in the approaches might be at the origin of the somewhat contrasting findings on the effect of financial constraints on a firm's exports. The identification assumption common to most measures is that they look at the credit supply as it is reflected by the perception or actions of the firm (Farre-Mensa and Ljungqvist, 2013). We instead build a proxy of financial constraints based on a credit rating index. This means that we identify where the firm credit supply becomes inelastic from the point of view of banks and credit institutions. By incorporating the credit markets' view and attitude towards potential borrowers, credit ratings measure the way investors decide to provide external finance. A similar approach based on credit scores is followed in Muuls (2008)'s study of Belgian firms, and in Wagner (2012)'s study of German firms.

Moreover, Italy represents a particularly interesting case to study. The industrial system is populated by a large number of small-medium sized firms, which are usually considered more exposed to financing problems. And, moreover, financing problems might be of particular relevance since the structure of the Italian financial system presents peculiarities with respect to other major countries. 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), bank credit plays a prominent role as a source of firm financing in Italy. In recent years, 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 (Panetta, 2013).[2] This is a persistent feature of the Italian system, with similar numbers found over the 2000-2003 period under study in this work, consistently across Italian regions (Bank of Italy, 2004). In this context, having a rating index that is issued by an agency "internal" to the banking system, and widely used by Italian banks, provides us with an ideal opportunity to disentangle the impact of financial constraints on exports.

Our findings confirm that constraints negatively affect exports, conditional on entry. However,

[2]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. Bond financing of Italian non-financial corporations amounts to less than 8% of firms' total financial debt.

the estimated effect of within-firm changes of financial status is sizably larger than the impact across constrained vs. unconstrained firms commonly identified in previous studies.

2 Empirical model and estimation strategy

Our main goal is to explore the relationship between financing constraints and a firm's exports. The baseline equation of interest is

$$\ln Exports_{f,t} = \gamma FC_{f,t-1} + \boldsymbol{\beta}\mathbf{Z}_{f,t-1} + FE_f + \epsilon_{f,t} \tag{1}$$

where $Exports$ is the value of exports of firm f in year t, FC_f is a dummy identifying constrained firms, $\mathbf{Z}_f$ is a set of firm level controls, FE_f is a firm fixed effect capturing time invariant firm-specific unobserved characteristics, and $\epsilon_{f,t}$ is a standard error term. The FC status and controls are lagged, as a first way to reduce potential simultaneity. The parameter of main interest is γ, capturing the effect of being financially constrained.

There are two potential sources of bias in estimating equation (1). A first issue is that, as suggested by economic theory and previous empirical evidence, firms self-select into exports. Thus, hidden factors affecting firms' decision to enter foreign markets likely correlate with unobserved factors influencing trade activities. Failing to account for this correlation may result into inconsistent estimates of the parameters of interest. Second, an endogeneity problem can arise from potential joint determination of export performance and financial constraints. Indeed, unobserved factors influencing the credit supply of firms might also influence firms' export values.

In order to cure for both potential sources of bias, and at the same time allowing for firm fixed effects, we apply the approach developed in Semykina and Wooldridge (2010). The method allows for consistent estimation of panel data models with selection also in presence of correlated unobserved effects and endogenous regressors.

The estimation strategy entails to explicitly model the selection mechanism via a two-equation, Heckman-type framework

$$\ln Exports_{f,t} = \gamma_1 FC_{f,t-1} + \boldsymbol{\beta}\mathbf{Z}_{f,t-1} + FE_{1f} + \epsilon_{1f,t} \tag{2}$$

$$s_{f,t} = 1\left[\gamma_2 IV^{FC}_{f,t-1} + \boldsymbol{\delta}_t \mathbf{W}_{f,t-1} + FE_{2f} + \epsilon_{2f,t} > 0\right] \quad . \tag{3}$$

Equation (2) is the main equation of interest (corresponding to Equation 1 above), where γ_1 is the coefficient of primary interest capturing the impact of the potentially endogenous FC dummy. Equation (3) is a Probit selection equation where $s_{f,t}$ is a binary indicator of a firm's export status (1 if a firm is exporter in t, 0 otherwise). Among the arguments of the indicator function $1[\cdot]$ on the right hand side, $IV^{FC}_{f,t-1}$ is the instrumental variable for $FC_{f,t-1}$, $\mathbf{W}_{f,t-1}$ contains exogenous explanatory variables, FE_{2f} is an unobserved firm fixed effect, and $\epsilon_{2f,t}$ a usual error term. Following Procedure 19.2 in Wooldridge (2010), the instrument $IV^{FC}_{f,t-1}$ is generated as the fitted probability from a Probit regression with the FC dummy as the dependent and taking as regressors an appropriate instrument capturing exogenous variation in financial constraints, plus the controls in $\mathbf{Z}_f$ and their time averages. Notice that $\mathbf{Z}_f \subset \mathbf{W}_f$, since the set $\mathbf{W}_f$ includes the same firm-level controls included in Equation (2), but it also needs to include a further exclusion restriction variable to cure selection bias.[3] This exclusion restriction variable must capture factors that influence the choice to entry into export markets, but unrelated to subsequent export performance. Proxies of sunk costs of exporting are traditionally accepted to meet this requirement, at least since Roberts and Tybout (1997).

Semykina and Wooldridge (2010) show that, because of the presence of firm-specific unobserved effects also in the selection equation (3), just adding the inverse Mills ratio and using a standard Fixed Effects estimator produce inconsistent estimates of Equation (2). However, a solution is available through adding time averages of all the exogenous explanatory variables both in the main equation (controls and generated instruments for FC) and in the selection equation (controls, generated instruments for FC, and proxy of sunk costs of exporting).[4]

3 Data and variables

The analysis exploits three datasets that we merge to obtain our working sample.

First, we have access to the Italian Foreign Trade Statistics (*Commercio Estero*, hereafter COE). This is collected by the Italian Statistical Office (ISTAT) and represents the official register of all trade flows involving Italian firms. It contains values (in thousands of euros) and quantities (in Kilos) of all export transactions by exported product and destination country. These are then aggregated to provide a firm level value of foreign sales, that we use as our dependent variable.[5]

Second, we use the Italian Register of Active Firms (*Archivio Statistico Imprese Attive*, ASIA), which is also maintained by ISTAT and covers the universe of Italian firms operating in all sectors of activity, 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).

Third, we access a firm level accounting dataset collected by the Italian Company Account Data Service (*Centrale dei Bilanci*, CB) and available through ISTAT. The CB dataset collects standard annual balance sheets and financial statements for all Italian *limited liability* firms. The long term institutional role of CB ensures high data quality, limiting measurement error. The CB dataset posits constraints to the analysis, since we only have access to data over the period 2000-2003, and we can only exploit a small subset of the firm annual reports, limiting the list of variables that we can employ

[3]This is solves the well known identification problems due to linearity of the inverse Mills ratio in Heckman-type estimators.

[4]More precisely, we model $FE_{2f} = \xi \bar{IV}_f^{FC} + \boldsymbol{\xi}\bar{\mathbf{W}}_f + a_{2f}$, where a bar indicates time averages of a variable, and we model $(a_{2f}|IV_f^{FC}, \mathbf{W}_f) \sim \text{Normal}(0, \sigma_a^2)$. This is equivalent to assume that FE_{2f} is related to IV_f^{FC} and to $\mathbf{W}_f$ only through their time averages, while the remainder is independent of IV_f^{FC} and $\mathbf{W}_f$. Likewise, the other implicit assumption is that in the main equation $FE_{1f} = \eta \bar{FC}_f + \boldsymbol{\eta}\bar{\mathbf{Z}}_f + a_{1f}$. This transformation, similar in spirit to Mundlak (1978), uses time averages of the explanatories computed over the entire sample of exporters and non-exporters and it is therefore free of selection bias (see Semykina and Wooldridge, 2010, for details).

[5]Only transactions involving very small values are left out of the COE data. According to ISTAT, firms in COE cover about 98% of trade flows (http://www.coeweb.istat.it/default.htm). Thus, firms which do not appear in the dataset are either marginal exporters or do not export at all. A detailed description of the requirements for a trade flow to be recorded in the case of Italy is in Bernard et al. (2015).

Table 1: Representativeness vis a vis total manufacturing

	Number of firms		Number of exporters		Export value (Bill. euros)	
Year	Universe	Our sample	Universe	Our sample	Universe	Our sample
2000	565,396	108,017	78,412	44,955	209.9	176.9
2001	560,657	111,749	79,577	46,364	223.8	192.9
2002	552,940	113,056	80,593	47,076	220.9	190.6
2003	541,835	112,441	79,356	46,492	218.1	182.9

Notes: number of firms, number of exporters and the total value of exports for the universe of manufacturing firms and for our sample.

in the estimates. At the same time, however, the CB dataset is the source for our proxy of financial constraints, and it is thus fundamental for our work.

After merging the three sources, we obtain a dataset that covers the entire population of Italian limited firms, over the period 2000-2003. The working sample is an unbalanced panel including a total of 149,362 firms active in manufacturing, of which 66,059 are exporters. In Table 1 we show that, compared to the total Italian manufacturing, our sample covers on average 20% of firms, about 58% of all manufacturers that do export, and 85% of the total value of exports. This is also a consequence of the fact that in Italy the legal status of limited firm is more common among medium-large firms, which are also those actors creating the vast majority of jobs, value added and exports. Indeed, in the years under analysis, limited firms represent about 65% of total manufacturing value added and account for about 75% of total employment. Notice, however, that we do have micro and small firms in the sample, and that we do not observe any strong under-representation of small or micro firms, as compared to their proportion in the population of Italian manufacturers.[6]

Measuring financial constraints

We base our assessment of firm-level financial constraints (FCs) on a firm-specific credit rating issued yearly by CB. This rating index results from an in-depth analysis conducted by professional financial analysts, complementing "hard data" on borrowers' annual reports with relevant soft information collected locally.[7] The index scores firms on a scale of 9 categories of decreasing 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.

The problem of measuring financial constraints represents a long debated issue. An accepted theoretical definition of constraints is a situation where a firm faces an inelastic credit supply schedule, so that the possibility to obtain external financial resources is ruled out. The key difficulty originates from the empirical impossibility to observe when this situation does happen in practice. To overcome

[6]The data were accessed at the ISTAT facilities in Rome, and are made available after careful screening to avoid disclosure of individual information. See Secchi et al. (2013) for more information on the data sources and their coverage.

[7]Soft information refers to any kind of information coming from private and direct firm-lender relationships, i.e. other than the relatively more public information about the availability of collateral or about administrative statements. See Petersen (2004) for a discussion of soft vs. hard information. A detailed description of the CB index is available at http://www.cervedgroup.com.

this problem, the literature on financial constraints, even outside trade studies, proposes a few indirect proxies of FCs. Scholars either rely on surveys, and thus define the constrained status based on what firms say and perceive, or resort to data coming from firms' financial accounts, defining constrained firms as those with, e.g., poor liquidity, high leverage, high cash-to-cash-flow sensitivity, or low collateral, on the presumption that these financial variables, or combination of the latter, strongly correlate with the ability to raise external credit. In reviewing the most common measures, Farre-Mensa and Ljungqvist (2013) explain that the implicit identification assumption common to the all the existing proxies is that "managers' opinions or actions reflect the shape of the credit supply curve as they perceive it."

Credit ratings represent a valid alternative to identify which firms are likely to face an inelastic credit supply: ratings indeed capture the shape of a firm's credit supply curve as the credit market perceives or, more precisely, defines it. Let us explain how the specific characteristics of the CB rating support this view.

First, the CB index is an official rating within the Italian banking system, extensively used by Italian banks in the evaluation of potential borrowers. This role of benchmark internal to the banking system is crucial to proxy for financial constraints dynamics in the Italian case, where bank credit is by far the primary source of external finance for firms, as we mentioned in the introduction.

Second, and relatedly, the role of the CB index as a key input in lending procedures of Italian banks makes it a close proxy for what banks actually do. This is confirmed by previous empirical analyses showing that there is a tight link between the CB rating and the availability and the cost of external finance. Guiso et al. (2013) provide clear evidence that, *ceteris paribus*, bad ratings have a clear association with higher interest rates and thus, with the cost of credit. Panetta et al. (2009) show that it is unlikely that a firm with poor rating can receive any credit.

Third, the CB index is a suitable measure of FCs also because of the complex set of information captured by the index itself, combining hard and soft data. The CB rating, in other words, does not merely work as a summary measure of firm performance, but it also includes other considerations which are taken into account in banks' lending decisions. Previous empirical analyses (Bottazzi et al., 2008, 2014) corroborate this idea, showing that an important fraction of highly productive, highly profitable and fast growing firms indeed receive very poor CB scores, so that there is not a trivial one-to-one mapping between performance and access to credit.

We exploit the CB rating index to split the sample into constrained and unconstrained firms. We build a financial constraints dummy (FC) that, in each different year, equals 1 if the CB rating of a firm is in category 8 or 9, and 0 otherwise.[8] As already noted in presenting the empirical model, we use the 1-year lagged value of the FC dummy in the regressions. Together with providing a first control for simultaneity, this choice is also appropriate because the rating scores are updated by CB at the end of each year, and it is therefore the rating in $t-1$ that is available to credit suppliers for their decisions about credit provision in year t.

[8]We label as NFC (Not-FC) firms those firms with FC dummy equal to 0. An alternative strategy could be to to use all (or to group some) of the original 9 rating classes, thus accounting for the graduation in the difficulty to access external finance. However, since the index is an ordinal variable, there is no quantitative meaning in moving, for instance, from class 4 to class 6. Our binary categorization, moreover, avoids the potential error in variables problem arising from including dummies for each of the rating classes.

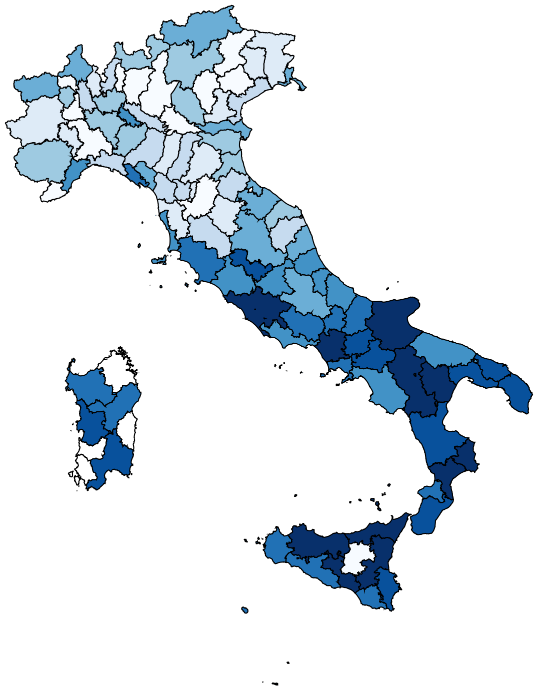

Figure 1: Geographical distribution of constrained firms, as share of FC firms over the total number of firms by province in 2003. All firms (left) and exporters (right). Darker areas reflect higher shares of FC firms.

According to our definition of FC firms, 17.5% of the firms in the data are financially constrained. The same ratio is 10% among exporters. There is also variability in firms' financial status over time. Indeed, looking at the 1-year transition matrix the percentage of firms moving from NFC to FC is 5.7%, while those moving from FC to NFC is 43.1%. Taking the longer time lag between t and $t+3$, the changes from NFC to FC status increase to 13.9% of total transitions, while changes from FC to NFC status amount to 40.6%. The percentage of within-firm changes in the FC status are even larger if we only consider the exporting firms separately. Between t and $t+1$ the share of firms moving from NFC to FC is 7.2%, while the FC-to-NFC changes account for 44.1% of total transitions. This is important for identification of the FC effect, since we want to exploit within-firm changes in FC status over time.

Figure 1 shows the share of FC firms by province, in the total sample (left) and among exporters (right).[9] The darker a province in the map and the higher the corresponding share of FC firms over the total number of firms in the province. We observe that, notwithstanding the relatively more limited presence of FC firms in Northern provinces, constrained firms are not clustered in few local areas: some Southern provinces do display low shares of rationed firms, as well as some Northern provinces see a relatively high presence of constrained firms. The picture is in line with figures in Minetti and Zhu (2011) where the definition of FCs exploits a survey of Italian firms asked to assess their perceived degree of credit rationing.

[9] A province is a local administrative entity, grouping several municipalities and cities, roughly corresponding to a US county.

Main controls

Guided by the literature on financing problems of firms (see Cabral and Mata, 2003; Almeida et al., 2004; Angelini and Generale, 2008, among others), and constrained by the limited set of CB variables available to us, we select a core of key firm-level attributes defining the set of controls $\mathbf{Z}$ in our main estimates. First, given the well established result that smaller firms tend to be more prone to financing problems, we control for firm size via the number of employees (*Empl*). Second, given that younger firms are known to have limited access to external financial resources, we include age (*Age*) computed from the year of foundation of each firm. Third, we include two proxies of financial factors that may interact with the provision of external finance in determining the overall amount of financial resources available to a firm. We use Gross Operating Margins (*GOM*) as a proxy of internal resources generated by operations, and the amount of total assets (*Assets*), that is including commercial credit, tangibles, intangibles and financial assets, which gives us a book measure of overall value of firms which creditors can pledge as "indirect collateral" in case of default or bankruptcy. All the controls enter the regression in logs and lagged by one year.[10]

Table 2 presents basic descriptives for the year 2003, distinguishing all vs. exporting firms (Panel A vs. Panel B).[11] In column 1, we confirm the stylized facts that exporters are on average bigger, older and that they have a stronger financial side, with more assets and more internal resources. Columns 2 and 3 break down the sample between constrained and unconstrained firms, and column 4 shows difference-in-means tests obtained by running an OLS regression of the FC dummy on firm attributes (in logs), including 3-digit industry fixed effects. Results obtained over the entire sample (in Panel A) confirm the stylized facts that firms affected by financial constraints tend, on average, to be smaller, younger, and to suffer from a relatively weaker financial structure in terms of less assets and less internally generated resources. The same picture is confirmed if conditioning upon being exporters (in Panel B). We also observe that constrained exporters export on average less than unconstrained exporters.

Instruments

Our econometric strategy requires a further set of variables to cure endogeneity of the FC dummy and self-selection into exports.

In order to build an instrument for the FC dummy, in the absence of firm level variables allowing to identify exogenous variation in firm level access to credit, we follow a common approach in the empirical literature on Italy, originally proposed in Guiso et al. (2004, 2006), that resorts to geographical variation of credit supply at the level of Italian provinces. The core idea is to look at the historical development of the regulation of the Italian banking system, and exploit the exogenous variation in the province-level distribution of credit due to the progressive removal, during the 1990s, of a series of restrictions to banking services. As explained in detail in Guiso et al. (2004, 2006), until the 1990s

[10]Export values and controls are deflated with appropriate sectoral price indexes available trough ISTAT at the 2-digit industry level, base year 2000.

[11]Given the panel is unbalanced, the number of firms is 112,441 in this year (see Table 1). Figures are stable in the other years.

Table 2: Descriptive statistics, 2003

	Our sample - Averages			Difference between FC and non-FC firms
	All firms	FC	non-FC	
	(1)	(2)	(3)	(4)
Panel A - All firms				
Number of employees	26.17	12.8	28.8	-0.863*** (0.011)
Age	14.56	10.1	15.4	-0.569*** (0.008)
Total Assets	5,539	2,969	6,053	-1.054*** (0.014)
Gross operating margin	493.1	72.8	577.5	-2.106*** (0.019)
Number of firms	112,441	18,735	93,706	
Panel B - Exporters				
Number of employees	48.82	33.2	50.5	-0.874*** (0.022)
Age	18.51	12.8	19.1	-0.644*** (0.015)
Total Assets	11,255	9,613	11,435	-0.868*** (0.025)
Gross operating margin	996.9	228.5	1,0810.9	-2.295*** (0.042)
Exports	3,775	2,980	3,862	-1.369*** (0.042)
Number of firms	46,492	4,580	41,912	

Notes: Columns 1-3 report, respectively, averages computed in 2003 over the total sample, and over FC or NFC firms, for all firms (Panel A) and for the subset of exporters (Panel B). Nominal variables in thousands of Euros, deflated. Column 4 shows difference-in-means tests between FC and NFC firms, again for 2003, performed over all firms (Panel A) and over exporters only (Panel B). The test is obtained via an OLS regression of the FC dummy on firm attributes (in logs), controlling for 3-digit industries. Robust standard errors in parenthesis. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

the distribution of banks and bank branches across Italian provinces came about in compliance with the rules implemented by the regulatory authorities in 1936. That regulation was fixing limits to the number of bank affiliates per province, in a way essentially unrelated to the structural characteristics and the level of economic development of the provinces themselves. The removal of these restrictions during the 1990s liberalized the credit market, opening up the possibility to establish new affiliates within and across provinces. This exogenous change had differentiated impact across provinces, also because the extent and strength of deregulation was differentiated depending on the type of banks involved (cooperative vs. saving banks, in particular) and that such different entities were unevenly spread across provinces at the time of deregulation.

Exploiting such historical patterns, we use the 1990-1999 difference in the number of bank branches (per 1,000 inhabitants) in each province as the key instrument to isolate exogenous variation of the FC dummy in our selectivity-endogeneity robust estimation (see below for details). Key for identification is to clarify the link between the chosen instrumental variable and the CB rating. In general, we expect more favorable credit ratings in provinces where we observe an higher number of 1990-1999 newly created bank branches. Indeed, a closer spatial proximity between firms and banks allow the latter to gather more and more precise information on firms, which should in turn get reflected in the soft data defining the CB index. This stands on two assumptions. First, an important

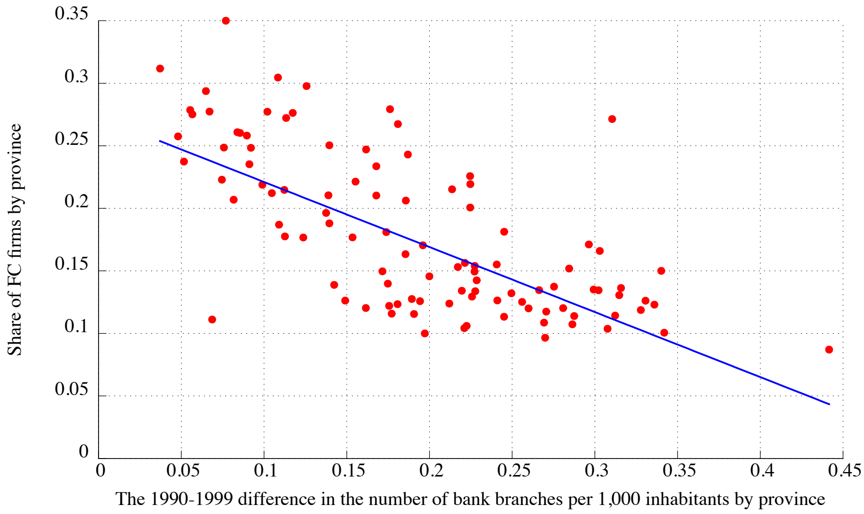

Figure 2: Relationship between the 1990-1999 difference in the number of bank branches and the share of FC firms by province, in 2003.

part of banks' knowledge of potential borrowing firms cannot be easily codified or transferred, and therefore it is only available locally, close to the place where a firm is located. Conversely, with less branches and thus less information, firms look more opaque, and the typical reaction by the banking system is a more conservative and distrustfully approach to evaluation of potential borrowers. This attitude, in turn, translates into attaching worse ratings to firms. Second, physical proximity between banks and firms, by itself, reduces opacity and increases the overall level of trust, for instance through repeated interactions. As a result, banks become more prone to accept potential borrowers' requests in local areas with a more widespread presence of bank branches, and this should translate into better rating scores by banks and other credit institutions.[12] Eventually, we expect the instrument to capture the variation in the ability of banks to collect more and more precise information on firms, in turn positively affecting credit ratings and thus reducing financial constraints. By contrast, the same variation is exogenous, in the sense that it is not expected to directly impact neither on firms' export behavior nor on unobserved firms' characteristics that determine export behavior.

Figure 2 provides empirical support to the validity of the instrument. It shows that the 1990-1999 difference in the number of branches in each province is indeed highly correlated with the share of FC firms in that province in 2003, with the expected negative sign. OLS and Least Absolute Deviation estimates of the slope of the relationship gives a coefficient of -0.516 (standard error 0.052) and -0.520 (standard error 0.067), respectively. Results are comparable in the other years in the sample period.

A further instrument required by the econometric methodology is a proxy for the sunk costs of exports, meeting the exclusion restriction that copes with selection bias. We define this proxy starting from the notion of Local Labour Systems (LLSs). These are geographical aggregations of municipalities defined by the Italian Statistical Office according to the degree of connectivity of the local labour market within the aggregations themselves, and thus identifying local areas where production-labour relationships are tight. Tight connections at the local level are likely characterized by activities such

[12] See Petersen and Rajan (1995); Bonaccorsi di Patti and Gobbi (2001); Carling and Lundberg (2005); Alessandrini et al. (2009) for further details on the role of distance on credit dynamics.

as sharing same trade services, accessing pools of established distribution networks, or exploiting neighbors' experience in dealing with foreign contracts and foreign legislation. These and possibly other factors tend to facilitate the entry into foreign markets, in turn reducing the sunk costs of exporting. Following Bernard and Jensen (2004) and Bernard et al. (2015), for each firm f, we define a proxy for the sunk cost of entry into exports ($ExpCost_f$) computed as the minimum between export entry and exit rates computed within the LLS wherein a firm is located. Higher rates of entry into or exit from export markets indicate lower sunk costs of exporting. There is substantial variation in the $ExpCost$ variable across LLSs and over time. In 2000, the median value of the variable is about 13% with a variance of 0.02, and it equals 9% and 21% for the 25th and 75th percentile, respectively. The over time variance across years within each LSS ranges from a minimum of 0.01 to a maximum of 0.25.[13]

4 Results

A consistent estimate of the FC dummy coefficient γ in our baseline regression (1) is obtained with the following steps:

Procedure 1

1. *generate the instrument $IV^{FC}_{f,t-1}$ as the fitted probability from a Probit regression of the binary indicator FC against the provincial level instrument for credit conditions (i.e. the 1990-1999 difference in the number of bank branches per 1,000 inhabitants in each province), the controls in $\mathbf{Z}_f$ (that is, the firm-level characteristics mentioned above) and their time averages;*
2. *obtain the inverse Mills ratio $\hat{\lambda}_{f,t}$ from a Probit estimate of equation* (3) *augmented with the time averages of the generated instrument $IV^{FC}_{f,t-1}$ from Step 1, and with time averages of the controls in $\mathbf{W}_f$ (that is, of the firm-level characteristics plus $ExpCost$);*
3. *estimate via pooled 2SLS-IV equation* (2) *augmented with the time averages of the generated instrument $IV^{FC}_{f,t-1}$, with the time averages of the explanatories in $\mathbf{Z}_f$, and with the inverse Mills ratio $\hat{\lambda}_{f,t}$ obtained in Step 2 together with its interactions with time dummies; use $\mathbf{Z}_f$, $IV^{FC}_{f,t}$, all the time averages and $\hat{\lambda}_{f,t}$ as instruments;*
4. *obtain analytic standard errors via the sandwich estimator provided in Semykina and Wooldridge (2010).*[14]

The estimator enjoys several nice features (Wooldridge, 2010): it is robust to mis-specification of the Probit model, it is more efficient than directly including the 1990-99 difference in the number of branches as an instrument into a usual 2SLS-IV procedure and, finally, it does not require to adjust the

[13] We use here the ISTAT definition of LLS for the year 2001, amounting to 683 areas.

[14] We generally report standard errors clustered at the firm level to allow for serial correlation of the error terms of a given firm. However, since the instrument used to generate $IV^{FC}_{f,t-1}$ vary at the province level, we have also re-run all the regressions clustering the standard errors at province level. Results are robust to this alternative treatment of the error terms.

Table 3: Financial Constraints and Total Exports - Main estimates

	$\ln Exports_{f,t}$ POLS (1)	$\ln Exports_{f,t}$ FE (2)	$\ln Exports_{f,t}$ Procedure 1 (3)	$\ln Exports_{f,t}$ POLS (4)	$\ln Exports_{f,t}$ FE (5)	$\ln Exports_{f,t}$ Procedure 1 (6)
$FC_{f,t-1}$	-0.134***	-0.091***	-1.674**	-0.131***	-0.102***	-0.763**
	(0.027)	(0.028)	(0.699)	(0.027)	(0.028)	(0.401)
$\ln Empl_{f,t-1}$	0.208***	0.124***	-0.031			
	(0.012)	(0.019)	(0.029)			
$\ln LP_{f,t-1}$				0.210***	0.028	-0.028
				(0.012)	(0.017)	(0.018)
$\ln Age_{f,t}$	-0.126***	-0.049	0.308***	-0.043	0.054	0.344***
	(0.012)	(0.075)	(0.095)	(0.012)	(0.074)	(0.095)
$\ln ASSETS_{f,t-1}$	0.947***	0.479***	0.392***	1.054***	0.567***	0.439***
	(0.011)	(0.031)	(0.028)	(0.007)	(0.031)	(0.029)
$\ln GOM_{f,t-1}$	0.071***	0.022***	-0.043	0.065***	0.024***	-0.005
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.018)
$\hat{\lambda}_{f,t}$			0.505***			0.213**
			(0.097)			(0.101)
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	No	Yes	Yes	No	Yes	Yes
R-squared	0.463	0.910	0.463	0.464	0.910	0.400
N.Observations	124,759	124,759	124,759	124,464	124,464	124,464

Notes: The dependent variable is a firm's value of exports (in log). $FC_{f,t-1}$ is a dummy for financially constrained firms. All the regressions include a constant term. POLS regressions in Columns 1 and 4 also include sector (3-digit) and province fixed effects. FE regressions are reported in Columns 2 and 5. Results from the Procedure 1 are reported in Columns 3 and 6. Robust standard errors in parenthesis below the coefficient, clustered at firm level. Asterisks denote significance levels (***:p<1%; **: p<5%; *: p<10%).

2SLS-IV standard errors. However, standard weak instrument diagnostics fail in this context (Nichols, 2007), and thus validation of instruments relies upon goodness of fit of the first step Probit.

Compared to previous studies, our explicit control for unobserved heterogeneity in both the selection and the primary equation gives additional confidence of proper identification of the key parameters. Estimation of the coefficient on the FC dummy indeed exploits variation within firm, over time, and thus allows to quantify the effect of within firm changes of FC status on within firm changes in export values.

Columns 1-3 of Table 3 report our main results from the baseline model. For completeness, in columns 1-2 we also show more standard pooled OLS (POLS) and Fixed Effects (FE) estimates. POLS identify the role of financial constraints by comparing constrained vs. unconstrained firms. The FE estimates look at within firm variation, but without controlling for selection and endogeneity. In both cases, results support that financial constraints associate with a reduction in foreign sales. The coefficient of the FC dummy in the FE specification is significantly smaller in absolute value than in the POLS estimates. This suggests a negative correlation between omitted variables and assignment to the FC class, as it is indeed expected for unmeasured factors such as managerial ability or quality, for instance.

In column 3 we present the more reliable estimates, where we address unobserved heterogeneity jointly with selection and endogeneity bias trough Procedure 1. The term $\hat{\lambda}_{f,t}$ is the inverse Mills ratio estimated in Step 2 of the Procedure: the strong statistical significance of the corresponding coefficient confirms that selection is indeed an issue. Also notice that the validity of the instrument for FC (i.e. the Probit fitted probabilities IV^{FC}) is supported in the preliminary Probit from Step 1 of the procedure (see Table 5 in Appendix): the coefficient on the 1990-1999 difference in the number of bank branches (per 1,000 inhabitants) in each province is statistically significant (p-value 0.002) and with the expected negative sign (-0.510).[15]

We confirm that financial constraints induce a significant reduction in the value of exports. The effect is sizable: the point estimate is -1.674 with a 95% confidence interval between -0.297 and -3.051, implying that financial constraints reduce exports by more than 35%, *ceteris paribus*. This implies a sensibly smaller effect than in Minetti and Zhu (2011) who find a point estimate of -3.043 with a 95% confidence interval between -0.489 and -5.597. Of course part of the difference comes from Minetti and Zhu (2011) failing to control for firm fixed effects, due to the cross-sectional nature of their data, in addition to differences in the definition of FC firms and in their coverage of a smaller Italian sample. Conversely, compared to other studies, our estimated effect is larger in magnitude, although a direct comparison cannot be established due to different measures of financial constraints and differences in the adopted empirical models. For Belgian firms Muuls (2008) observes that a one unit increase in the credit score (i.e. a reduction in financial constraints in her case) increases exports by 8.8%. Li and Yu (2009) find that the elasticity of exports of Chinese firms to interest expenses is close to unitary. Damijan et al. (2010) analysis of Slovenian firms reveals that the coefficient between lagged debt-to-asset ratio and export intensity is positive and significant for both small and large firms, and equal to 0.054 and 0.142, respectively. Finally, Wagner (2012) notices that his estimates of a fractional logit model, exploiting a credit score as a measure of FCs for German firms, cannot be interpreted in a straightforward way in terms of elasticity or effect of unit changes on export values.

It is also remarkable that the hampering effect of FCs is stronger than what we can infer from the simple FE estimates in column 2. The latter are upward biased (smaller FC coefficient in absolute value), suggesting that the endogenous component of our FC classification associates with an underestimation of the true detrimental effect of being constrained on exporting activities. Concerning the controls, and limiting to the more reliable selection-endogeneity corrected estimates in column 3, we find that age and assets have a positive and strongly significant coefficient. The elasticity of exports to size and internal resources are not significant.[16]

In columns 4-6 we replicate the estimates including labour productivity, which is a potentially key omitted variable, as suggested by the crucial role that productivity plays in the literature on firms' heterogeneity in international trade. In line with the notion (essentially based on the Melitz's single-input framework) that size and productivity maps one-to-one, we report results of a model where we include labour productivity and exclude size (employment). Our main conclusion about the effect of FCs remains valid: the point estimate in column 6 goes down as compared to column 3, but it is equal to the baseline estimate if we take a 1-standard error confidence band. We thus keep employment in

[15]See Secchi et al. (2013) for further analysis of the goodness of fit of the first step Probit, confirming the validity of the instrument.

[16]We do not over-emphasize these results, however, since endogeneity of controls might be an issue. We tried a panel-GMM specification (also including lagged exports) to cure for that, but the short time span available did not allow to identify a set of lagged variables that passes standard Sargan-Hansen tests of exogeneity. Moreover, standard GMM methods do not allow to control for selection bias.

the rest of the paper.[17]

Robustness checks

We explore the robustness of the negative effect of FCs on exports through a series of additional exercises. Table 4 presents the results. All the reported estimates are obtained through application of Procedure 1, and are therefore informative about within firm changes of FC status corrected for selection and endogeneity.[18]

First, in column 1 we check robustness with respect to possible outliers, by removing the top and bottom 1% of the observations in all variables. Results are basically identical to the main estimates.

Second, in column 2 we use sales in place of employment as a proxy for size. This helps to focus on the relative importance of foreign sales over total sales, rather than simply controlling for the scale of operation via employment figures. Results remain practically unchanged.[19]

Third, we enrich our baseline regression model to include a firm-level proxy of costs among the regressors. If financial constraints are associated to higher costs, possibly due to inefficiency of constrained firms, then one wants to disentangle this aspect from a pure FCs effect. Seeking to capture this cost-efficiency channel, we include the (log of) the cost of labour per employee (*LabCosts*), obtained as the ratio between total labour expenses (wages and compensation) and number of employees. The results in column 3 confirm that financial constraints reduce a firm's exports. The point estimate of the FC coefficient is smaller (-1.499), but it is not statistically different from the baseline estimate in Table 3 within a 1-standard error confidence interval. The coefficient on labour cost is negative and significant: more costly labour reduces exports, confirming the cost-efficiency interpretation of this variable. Also notice that, compared to the baseline results, the inclusion of labour costs makes the "selection term" not statistically significant ($\hat{\lambda}_{f,t} = 0.013$), suggesting that the additional regressor absorbs self-selection into exports.

Fourth, we exclude from the analysis firms that start exporting during the time-span considered. Indeed, credit constraints might also affect entry into export markets, and a bias can arise if more recent export market entrants are more credit constrained than incumbent exporters. In column 4, we only consider continuous exporters, and use as the control group those firms that never serve foreign markets within the sample period. Following Greenaway et al. (2007) we define as "continuous exporters" those firms that export in all the years in which they are present in the sample, while "never exporters" are firms that never export within the sample period. The estimated FC coefficient

[17]We have also experimented with including a TFP measure of productivity. This was computed via the GMM modified Levinsohn-Petrin production function estimation proposed in Wooldridge (2009), with value added as output proxy, employment and fixed tangible assets as labour and capital inputs, respectively, and materials as instrument. The estimated FC coefficient turned out as not statistically different from our baseline estimates. We do not use TFP further in the study, however, since the exercise suffers from imprecise GMM estimation of production functions due to the short time period available, and from problems with measurement of the capital proxy. In the CB data, indeed, fixed tangible assets present several zeros or missing values, considerably reducing the sample, and they are reported as gross of depreciation, but we do not have investment figures needed to recover the net value through the perpetual inventory method.

[18]Note that for each robustness check, the Probit used to build the instrument $IV^{FC}_{f,t-1}$ is correspondingly adapted. Results and related goodness of fit tests are available upon request.

[19]Sales however create collinearity problems, with total assets in particular, as revealed by standard Variance Inflation Factors test from OLS estimates of our main equation. This also motivates our use of employment as main proxy for size.

Table 4: Firm Financial Constraints and Total Exports - Robustness

	$\ln Exports_{f,t}$ Outliers (1)	$\ln Exports_{f,t}$ Sales (2)	$\ln Exports_{f,t}$ Lab.Cost (3)	$\ln Exports_{f,t}$ No Entry (4)	$\ln Exports_{f,t}$ No MNCs (5)	$\ln Exports_{f,t}$ Different FC (6)	$\ln Exports_{f,t}$ Province Controls (7)	$\ln Exports_{f,t}$ Initial Development (8)
$FC_{f,t-1}$	-1.966***	-1.548**	-1.499**	-1.987***	-1.433**	-1.554**	-1.418**	-1.656**
	(0.770)	(0.663)	(0.642)	(0.452)	(0.627)	(0.699)	(0.670)	(0.691)
$\ln Empl_{f,t-1}$	-0.043		-0.031	0.071***	0.007	-0.027	-0.026	-0.031
	(0.030)		(0.034)	(0.022)	(0.026)	(0.030)	(0.029)	(0.029)
$\ln Age_{f,t}$	0.273***	0.467***	0.247**	-0.105	0.233**	0.406***	0.334***	0.308***
	(0.096)	(0.098)	(0.099)	(0.092)	(0.100)	(0.103)	(0.094)	(0.095)
$\ln ASSETS_{f,t-1}$	0.377***	0.489***	0.589***	0.493***	0.591***	0.497***	0.388***	0.392***
	(0.029)	(0.051)	(0.037)	(0.026)	(0.035)	(0.027)	(0.027)	(0.028)
$\ln GOM_{f,t-1}$	-0.057	-0.043*	-0.034	-0.048***	-0.031	-0.053	-0.032	-0.042
	(0.033)	(0.024)	(0.026)	(0.017)	(0.025)	(0.027)	(0.028)	(0.029)
$\ln Sales_{f,t-1}$		0.128**						
		(0.055)						
$\ln LabCosts_{f,t-1}$			-0.069**					
			(0.028)					
$\hat{\lambda}_{f,t}$	0.453***	0.136*	0.013	0.367***	0.243**	0.435***	0.420***	0.508***
	(0.159)	(0.067)	(0.098)	(0.066)	(0.109)	(0.097)	(0.091)	(0.097)
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.385	0.387	0.388	0.400	0. 369	0.387	0.397	0.388
N.Observations	118,357	124,472	120,352	99,459	115,735	124,759	124,759	124,759

Notes: The dependent variable is firm value of exports (in log). $FC_{f,t-1}$ is a dummy for financially constrained firms. All the regressions include a constant term and are estimated via Procedure 1. Column 1: outliers (top and bottom 1% of the variables distributions) are removed. Column 2: sales as size in place of employment. Column 3: labour costs per employee is included. Column 4: only continuous and never exporters. Column 5: multinational firms are excluded. Column 6: different definition of the FC dummy, including firms rated as 7. Column 7: additional province-level controls are added, i.e. (log of) value added, (log of) population and an index for the level of infrastructure. Column 8: different instrument, i.e. the 1990-1999 difference in the number of branches (per 1,000 inhabitants) divided by the number of branches (per 1,000 inhabitants) in 1990. Robust standard errors in parenthesis below the coefficients, clustered at firm level. Asterisks denote significance levels (***:p<1%; **: p<5%; *:p<10%).

is larger, but not statistically different from the baseline estimate within a 1-standard error confidence band.

Fifth, in column 5, we re-estimate the model after excluding all multinational companies (MNCs), since this type of firms are known to have quite specific export dynamics. Our baseline findings are not affected: the point estimate of γ_1 is smaller, but equal to the full sample results within a 1 standard error confidence band.

Next, we test the robustness with respect to a different definition of the FC group. In column 6, we include in the FC class also the firms with a CB rating of 7. This means a less restrictive threshold to the identification of inelastic credit supply. We observe a reduction in the estimated FC dummy coefficient as compared to the baseline results, as expected given the milder definition of the FC group. Once again, however, the estimated FC effect is not statistically different from the baseline estimates within 1 standard error confidence interval.

Finally, in columns 7-8, we address potential weaknesses affecting the instrument for FC. First, a potential violation of the exclusion restriction on the instrument may arise from correlation between the instrument itself and idiosyncratic province-level components of the error term that are in turn related to firms' exports and to the level of development of the banking sector. To check for this potential problem, we re-estimate our baseline model including additional province-level variables.[20] In column 7, 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 (jointly produced by the Association of Italian Chambers of Commerce in collaboration with the "Guglielmo Tagliacarne" Institute). The main conclusion is not affected: the estimated effect of financial constraints is still negative and significant, and it does not differ from the baseline estimates within a 1-standard error band. Further, in column 8, we control for the level of financial development of the province at the beginning of the deregulation period, by normalizing the instrument for the number of branches in 1990. Also in this case the main result of a negative and significant FC effect is preserved, and the magnitude is statistically comparable to the baseline regression.

5 Concluding Remarks

In this paper we exploit information on credit ratings internal to the Italian banking system to re-examine the causal effect of firm-level financial constraints on the export performance of a large panel of Italian firms. Our empirical strategy allows to control for unobserved firm characteristics both in the selection and in the main equation, at the same time curing potential endogeneity of financial constraints via an instrumental variable approach. The few previous attempts to get rid of self-selection into exports and endogeneity fail to control for firm heterogeneity. This implies that, net of selection and endogeneity bias, they can only assess differences in export performance across constrained vs. unconstrained firms, while our article is the first estimating the effect of a within-firm change in credit conditions on a firm's exports. As a result, we can be more confident that our findings provide a valid contribution to the debate on the importance of credit constraints for the performance

[20]Including province fixed effects does not work, since IV identification exploits variation across provinces.

of exporters.

Our key result is that, conditional on entry into international markets, financial constraints cause firms to export less in value. This within-firm, selection-endogeneity corrected effect of financial constraints is large, in general terms. And, in particular, it is larger than the FCs effect we obtain from pooled OLS estimates comparing constrained vs. unconstrained firms, and also larger than the within-firm effect obtained without correcting for selection and endogeneity. Moreover, it tends to be larger than what has been suggested by most previous studies, although different samples, countries and empirical methodologies do not allow for precise comparisons.

The estimated contraction in the intensive margin due to financing problems corroborates theoretical predictions that financial constraints do not only matter to cover costs related to entering foreign markets in the first place, but they are also relevant to cover variable costs of exporting. The finding has also relevant policy implications, particularly to draw lessons for the understanding of the possible consequences of the recent crisis. Indeed, it shows that the credit contraction may favor a process of selection based on firms' deep pockets, rather than "rewarding" productivity, with negative effects on the overall competitiveness of the economic system. Given the large magnitude of the estimated effect of financing constraints from our study, a careful evaluation of the complex interactions between firm heterogeneity, financial frictions, and aggregate dynamics appears crucial for designing appropriate policy intervention.

Appendix

In implementing Procedure 1, we follow Wooldridge (2010) and use as instrument the fitted probability $IV^{FC}_{f,t-1}$ from a Probit of our binary indicator FC on: (a) the provincial level instrument capturing exogenous variation in credit conditions (i.e. the 1990-1999 difference in the number of bank branches per 1,000 inhabitants in each province); (b) on the firm-level controls in $\mathbf{Z}_f$; and (c) on their time averages.

Table 5 reports the results of this Probit. As expected, the coefficient on the FC dummy is negative and strongly significant, confirming the validity of the instrument. Controls have also the expected negative sign, when significant.

Table 5: First Step Probit Estimates

	$FC_{f,t-1}$
1990-1999 Difference in # of bank branches	-0.510***
	(0.162)
$\ln Empl_{f,t-1}$	-0.048***
	(0.012)
$\ln Age_{f,t}$	0.067
	(0.075)
$\ln ASSETS_{f,t-1}$	-0.091***
	(0.013)
$\ln GOM_{f,t-1}$	-0.014***
	(0.003)
Year FE	Yes
Firm FE	Yes
Pseudo R-squared	0.209
N.Observations	274,181

Notes: The dependent variable is the FC dummy. Robust standard errors in parenthesis below the coefficients, clustered at province level. Asterisks denote significance levels (***:p<1%; **: p<5%; *:p<10%).

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