DISCUSSION PAPER SERIES

No. 7975

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INTERNATIONAL TRADE AND REGIONAL ECONOMICS

Centre for Economic Policy Research

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ISSN 0265-8003

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Holger Breinlich, CEPR and University of Essex
Alessandra Tucci, European Commission Directorate General for Economic and Financial Affairs

Discussion Paper No. 7975
September 2010

Centre for Economic Policy Research
53–56 Gt Sutton St, London EC1V 0DG, UK
Tel: (44 20) 7183 8801, Fax: (44 20) 7183 8820
Email: cepr@cepr.org, Website: www.cepr.org

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ABSTRACT

Foreign Market Conditions and Export Performance: Does 'Crowdedness' Reduce Exports?*

This paper analyzes the link between firm exports and the competitive environment in foreign markets. We derive a theory-based econometric specification linking destination-specific exports to foreign demand and the degree of 'crowdedness' of foreign markets. The latter is a measure of the number and efficiency of firms competing in a given market and the barriers impeding their access.

We estimate this specification on a large sample of Italian manufacturing firms between 1992 and 2003 and use the results for counterfactual experiments. We find that increases in the crowdedness of foreign markets have reduced Italian exports by around 0.2%-0.3% per year. However, other factors such as higher unit labor costs and weak demand growth in Italy's main export market (the EU15) have been much more important in explaining Italian export performance. Our results also indicate that China's impact on Italian exports is small and if anything positive.

JEL Classification: F12, F13 and F15
Keywords: competition, exporters, foreign markets and international trade

Holger Breinlich
Department of Economics
University of Essex
Wivenhoe Park,
Colchester CO4 3SQ
United Kingdom

Email: hbrein@essex.ac.uk

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Alessandra Tucci
European Commission,
DG Economic and Financial Affairs,
BU-1 3/179,
B-1049 Bruxelles
BELGIUM

Email: a.tucci@lse.ac.uk

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* We would like to thank Joao Santos Silva, Frédéric Robert-Nicoud, Werner Antweiler and two anonymous referees for helpful comments and suggestions. This paper has also benefited from the comments of seminar and conference participants at the Universities of Essex, Warwick and Stockholm, the Stockholm School of Economics, Royal Holloway, Sussex, ETSG 2007, EIIE 2008, the EER Young Economist Workshop 2008 and the VfSP Meeting 2008. Gunes Asik-Altintas provided competent research assistance. We gratefully acknowledge financial support through British Academy grant SG-47030. This paper reflects the views of the authors only and should not be attributed to the European Commission. This paper was prepared and submitted for publication while Alessandra Tucci was at the Centre for Economic Performance at the London School of Economics, Houghton Street, London WC2A 2AE.

Submitted 16 August 2010

1 Introduction

Exports make up a large and growing share of total manufacturing sales in most industrialized economies. For example, the ratio of total manufacturing exports to production in 2003 was 16% in the USA, 42% in the UK, 47% in Germany and over 70% in small open economies such as Belgium, Denmark or the Netherlands. For all OECD countries, this ratio was on average 53% in 2003, up from 35% in 1990 and 24% in 1970 (OECD, 2005). At the same time, there have been growing concerns in industrialized countries about the rise of large emerging economies – Brazil, India and especially China – and the 'threat' this poses to domestic exporters. Given the increasing importance of foreign markets for manufacturing sales, what impact will these changes have on the export performance of developed economies? More generally, how do competitive conditions on foreign markets affect exports of domestic firms? Are these conditions key determinants of export success or are other factors such as foreign demand or firm-level unit labor costs more important?

In this paper, we use a large dataset on Italian manufacturing firms to address these questions. We employ a firm-level gravity model to derive an econometric specification linking destination-specific exports to firm characteristics, foreign demand and the competition intensity or "crowdedness" of foreign markets.[1] This latter variable will be at the centre of our analysis. It is a measure of the number and efficiency of firms competing in a given market and the barriers impeding their access, such as tariffs or physical distance. It summarizes how easy or difficult it will be for an exporter to penetrate a given market, holding constant other factors such as foreign demand or the unit labor costs of the exporting firm. The principal goal of this paper is to quantify the role of market crowding and its components and to compare their quantitative importance to other determinants of export performance.

We proceed in three steps. Having derived our econometric specification, we estimate it on a large sample of Italian manufacturing firms between 1992 and 2003. We find that market crowding has a robustly negative impact on firm-level exports across a wide range of specifications and that its effect operates both along the extensive and the intensive margin. We also show that the same does not hold true for a number of alternative, non-structural measures of foreign competition intensity.

We then examine the quantitative importance of our findings more closely by performing a series of counterfactual experiments. Our findings indicate that increased numbers and efficiency of for-

[1] The New Economic Geography literature also uses the term "market crowding". We use these expressions in the rest of the paper since – as will become clear below – our measure is somewhat different from the standard usage of the word "competition intensity" in industrial organization.

eign firms, and improvements in their access to destination markets, have reduced Italian exports by around 0.2%-0.3% per year. This is similar in size to the effects of tariff reductions for Italian firms (+0.3%/year) but smaller than the impact of higher unit labor costs (-1.4%/year) and less favorable exchange rates (-2.0%/year). By far the most important determinant of export performance was foreign demand growth, however, raising Italian exports by up to 5% per year or 55% over the sample period. Our results also indicate that China's overall impact on Italian export performance is small and if anything positive at around +0.2%/year. Much more important in explaining the slow growth of Italian firms' exports has been the relatively slow demand growth in Italy's main export market, the EU15.

We believe that these findings are important for a number of reasons. From a policy perspective, Italy is an interesting case to study since its exporters have been losing world market shares for over a decade. This is often linked in public debates to the emergence of competitors from low-wage countries like China which compete head to head in traditional Italian export sectors such as apparel or textiles. Our finding that the increased crowdedness of foreign markets is not the principal determinant of Italian export performance sheds doubt on this conjecture.

Our findings also contribute to the wider issue of firm-level responses to trade integration. The traditional focus of this literature has been on the effects of import penetration on a firm's home market, particularly in the wake of trade liberalizations (see Pavcnik, 2002, and Trefler, 2004, for two recent influential contributions; Tybout, 2003, provides a survey of the earlier literature). In contrast, our analysis quantifies – among other things – the effects of lower trade barriers on *foreign* markets. While the two issues are evidently related, there are also important differences. First, exporting firms are usually quite different from purely domestic firms. As previous research has shown, exporters tend to be larger, more productive, use more capital intensive production and employ a more highly skilled workforce (see for example Bernard and Jensen, 1995 and 1999; Wagner, 2007, and Greenaway and Kneller, 2007, provide surveys of the literature). Secondly, exporters will have more options at their disposal to react to increased market crowding than purely domestic firms – for example, redirecting exports to less crowded markets. On the other hand, the set of potential intervention mechanisms available to policy makers is more limited. This is because traditional instruments for protecting domestic firms from import penetration (tariffs, quotas) are evidently not available to national governments in this new setting.[2] Taken together, these considerations suggest that the reaction of exporters to changes on foreign markets might be quite different from the reactions of

[2] A corollary to this is that the usual econometric problems associated with the potential endogeneity of domestic tariffs and other forms of protection (see Trefler, 2004, for example) are likely to be less relevant in our setting.

domestic firms to increased import penetration which have been studied so far.

From a methodological point of view, our empirical measure of market crowding provides a new way of analyzing what Anderson and van Wincoop (2003) label “multilateral resistance”. As these authors explain, controlling for the multilateral resistance (or crowdedness) of a market is necessary to obtain consistent parameter estimates in gravity equation estimations. We go beyond simply controlling for multilateral resistance and decompose it into its different components – number and efficiency of competitors and the barriers impeding their access (tariffs, distance etc.).

Our analysis is also related to contributions by Redding and Venables (2003) and Hanson and Robertson (2006). These authors use gravity models to decompose changes in South-East Asian and Mexican exports, respectively, into contributions of the supply characteristics of the exporting countries and foreign market conditions. Using a similar methodology, Hanson and Robertson (2008) analyze the impact of changes in Chinese supply capacity on the exports of ten developing countries. All three contributions rely on country- or sector-level trade data. This precludes any analysis of the extensive and intensive margin of firm exports, which has featured prominently in the recent literature (e.g., Helpman et al., 2008). It also prevents these authors from looking at how the impact of foreign market conditions varies across firms, which is an important part of our analysis. Finally, as we argue in more detail below, the use of firm-level data makes it less likely that our results suffer from reverse causality problems, given that each individual firm only accounts for a small share of any particular foreign market. This is less likely to be true when looking at entire sectors or even countries.

A final paper related to our analysis is Bernard and Jensen (2003). These authors regress growth rates of U.S. firm-level exports in 1987-1992 on exchange rate variations, firm productivity and a measure of foreign income. They do not analyze the role of export market crowding and their data do not allow a destination specific analysis.

The rest of this paper is organized as follows. Section 2 presents a model of firm-level export behavior and introduces our empirical measure of market crowding. Section 3 describes the data and Section 4 presents the econometric results. Section 5 uses our estimates for various counterfactual experiments with regards to Italian firm-level exports. Section 6 concludes.

2 Theoretical Framework

We base our empirical analysis on a partial equilibrium model of firm exports in which firms face CES demand and operate under monopolistic competition. This framework is the workhorse of most of current international trade theory and has a number of advantages over possible alternatives, both in terms of predictive power and analytical convenience.

Most importantly, CES generates a log-linear specification relating export demand to importer and exporter characteristics and bilateral trade costs. As a vast empirical literature on gravity equation estimation has shown, this specification provides an excellent fit to international trade data at different levels of aggregation (see Anderson and van Wincoop, 2004, and Disdier and Head, 2008, for recent overviews). Our framework also has the advantage of comparability with existing theoretical and empirical work which mostly also builds on similar frameworks (e.g., Anderson and van Wincoop, 2003; Melitz, 2003; Helpman et al., 2008). Finally, CES allows to conveniently summarize the degree of market crowding in a single measure, the CES price index.

2.1 Firm-level exports

Assume that consumers in market n have identical CES preferences over the different varieties produced by firms in sector s. The demand facing any firm i in this sector from market n then takes the form $d_{ins} = p_{ins}^{-\sigma_s} P_{ns}^{\sigma_s - 1} E_{ns}$, where p_{ins} is the c.i.f. price charged by the firm in market n, E_{ns} is total industry-specific expenditure in market n and $\sigma_s > 1$ denotes the elasticity of substitution between varieties in industry s. $P_{ns} = \left(\sum_j \int_{i_{jns}} p_{ins}^{1-\sigma_s} di \right)^{\frac{1}{1-\sigma_s}}$ is the CES price index which measures the degree of crowdedness in market n, sector s. The index j denotes all countries exporting to n while i_{jns} denotes the exporters from each of these countries. In our data, each firm is classified into a single industry, so from now on we index firm-specific variables by the subscript i only.

In order to enter foreign markets, firms have to make upfront investments such as adapting products to local standards or setting up distribution channels (see Roberts and Tybout, 1997; Bernard and Jensen, 2004). The costs of doing so are equal to F_{in}. Firms also incur variable trade costs when exporting. These are $\tau_{in} - 1$ in terms of the exported good for each unit shipped to market n. Finally, revenues from market n have to be converted back to the home market's currency at the exchange rate e_{in}, expressed in units of the home currency per foreign currency unit.

With monopolistic competition and CES preferences, firms set prices at a constant markup over marginal costs:[3]

[3] We also experimented with alternative frameworks allowing for variable price-cost margins (e.g., Ottaviano and Melitz, 2008). However, our results indicated that the absence of income effects and the linearity of the resulting demand functions makes the Ottaviano and Melitz framework less suitable for empirical work on firm-level exports (the fit of our regressions was substantially lower, indicating that the functional form implied by Ottaviano and Melitz does not capture the data generating process well). In any case, our empirical proxy for the CES price index will be more general than its theoretical counterpart. Its components will capture both the direct effect of market crowding on firm-level demand (present in the model) and the indirect effect via reduced price-cost margins (absent from our model).

$$p_{in} = \frac{\sigma_s}{\sigma_s - 1} \tau_{in} c_i e_{in}^{-1}$$

We assume that the marginal costs of production, c_i, are constant. The choice of export price and quantity in market n is thus independent of the situation on other markets. With this pricing rule, the value of exports by firm i to market n is

$$r_{in} = p_{in} d_{in} = \left(\frac{\sigma_s}{\sigma_s - 1} \right)^{1-\sigma_s} \tau_{in}^{1-\sigma_s} e_{in}^{\sigma_s - 1} c_i^{1-\sigma_s} P_{ns}^{\sigma_s - 1} E_{ns} \tag{1}$$

and the price index can be expressed as

$$P_{ns} = \left(\frac{\sigma_s}{\sigma_s - 1} \right) \left(\sum_j \tau_{jns}^{1-\sigma_s} e_{jn}^{\sigma_s - 1} n_{jns} \frac{\int_{i_{jns}} c_{i_{jns}}^{1-\sigma_s} di}{n_{jns}} \right)^{1/(1-\sigma_s)} \tag{2}$$

where n_{jns} is the number of firms from j exporting to market n in sector s.

Note that firms will only export if the variable profits from doing so are at least equal to the initial setup costs F_{in}. Noting that variable profits are $\pi_{in} = \frac{e_{in} r_{in}}{\sigma_s}$, we obtain a market entry condition for firm i in terms of its marginal costs, setup costs F_{in}, market specific characteristics and bilateral trade costs. That is, firm i will enter a market n if and only if:

$$D_{in} \equiv \left(\frac{e_{in}^{\sigma_s} (\sigma_s - 1)^{\sigma_s - 1} E_{ns} P_{ns}^{\sigma_s - 1}}{c_{in}^{\sigma_s - 1} F_{in} \tau_{in}^{\sigma_s - 1} \sigma_s^{\sigma_s}} \right)^{1/(\sigma_s - 1)} \geq 1 \tag{3}$$

Expressions (1) and (3) form the basis of our econometric specifications. We can summarize a firm's export decision as

$$r_{in} = \begin{cases} \left(\frac{\sigma_s}{\sigma_s - 1} \right)^{1-\sigma_s} \tau_{in}^{1-\sigma_s} e_{in}^{\sigma_s - 1} c_i^{1-\sigma_s} P_{ns}^{\sigma_s - 1} E_{ns} \text{ if } D_{in} \geq 1 \\ 0 \text{ otherwise} \end{cases} \tag{4}$$

To reiterate, by estimating (4) we perform a partial equilibrium analysis, taking the number of competitiors and their prices, exchange rates, as well as foreign demand as given. In the full general equilibrium of our model, these will be determined as a function of underlying preference and technology parameters. We believe that a partial equilibrium approach is better suited here, since finding empirical proxies for the right-hand side elements of (4) is relatively straightforward – which is not true for the underlying parameters determining them.[4] A direct econometric implication of

[4] For example, we experimented with a version of Chaney (2008) which would have required proxies for the number of potential entrants in a market, market-sector specific fixed entry costs and market-sector specific cutoffs of the underlying

the partial equilibrium nature of our analysis is that we have to assume that individual Italian firms' influence on the destination-specific variables in (4) is negligible. Given that the average share of firms in our sample in the total sales volume of foreign markets is less than 0.0025%, we believe that reverse causality issues are indeed unlikely and this assumption thus justifiable.[5] There are of course additional endogeneity concerns arising from omitted variable bias. We address these in a number of ways in our empirical analysis below.

2.2 Choice of empirical proxies

We now turn to the choice of empirical proxies for the variables in (1) and (3).

Market Crowding - CES Price Index An empirical proxy for the price index P_{ns} requires data on $\tau_{jns}^{1-\sigma_s}$, $e_{jn}^{\sigma_s-1}$, n_{jns}, and $n_{jns}^{-1}\int_{i_{jns}} c_{i_{jns}}^{1-\sigma_s} di$. Exchange rate data are easily obtainable. While our theoretical model features full exchange rate pass-through, we want to allow for a less than perfect pass-through in the empirical analysis. We thus proxy $e_{jn} = \mu_1 ex_{jn}^{\alpha_1}$ where ex_{jn} denotes the bilateral exchange rate between j and n, and μ_1 and α_1 are parameters to be estimated below.

We do not have internationally comparable data on the number of exporters (n_{jns}) for all countries j appearing in P_{ns} (see Section 3 for a description of our data). We thus write the number of exporters n_{jns} as a function of the number of establishments in country j, sector s (est_{js}), multiplied by the share of n in country j's exports (share_{jn}). That is, $n_{jns} = \mu_2 (\text{est}_{js} \times \text{share}_{jn})^{\alpha_2} \equiv \mu_2 \nu_{jns}^{\alpha_2}$, where again μ_2 and α_2 allow for a more flexible functional form. The intuition underlying this approach is that the number of establishments in sector j provides a natural upper bound to the number of exporters in that sector. Multiplying this number by the share of foreign market n in j's exports reflects the empirical regularity observed by Eaton et al. (2004) – and present in our data as well – that a larger fraction of domestic firms exports to larger destination markets. In this sense, est_{js} and share_{jn} are both theoretically meaningful components of our exporter proxy. In the technical appendix to this paper (Section 3.1), we present additional results showing that using est_{js} as part of the proxy also improves the robustness of our crowdedness measure. Finally, as we explain in more detail below, the use of aggregate rather than sectoral export shares (i.e., share_{jn} rather than share_{jns}) will facilitate the estimation of the parameters μ_2 and α_2 through a gravity equation approach while still using the information contained in trade flows to help proxy for the number of exporters.

We also do not observe individual firms' marginal costs (c_i). We thus proxy $n_{jns}^{-1}\int_{i_{jns}} c_{i_{jns}}^{1-\sigma_s} di$ by

productivity distribution of potential entrants (assumed to be Pareto by Chaney).

[5] Even for the EU15, Italy's main export market, the average firm's market share is just 0.004%.

the average unit labor costs (the total wage bill divided by value added) in sector s, country j. That is, $n_{jns}^{-1}\int_{i_{jns}} c_{i_{jns}}^{1-\sigma_s} di = \mu_3 \left(uc_{js}\right)^{\alpha_3(1-\sigma_s)}$. This captures the intuition that the presence of firms from countries with lower production costs will make a given export market a tougher place to sell to. Again, the inclusion of the parameters μ_3 and α_3 increases the degree of flexibility of this functional form.

Third, we write trade costs as a log-linear function of variables commonly used in gravity equation estimations

$$\tau_{jns} = \mu_4 \text{dist}_{jn}^{\alpha_4} \times \mu_5(1+t_{jns})^{\alpha_5} \times \mu_6 e^{\alpha_6 \text{lang}_{jn}} \times \mu_7 e^{\alpha_7 \text{int}_{jn}} \tag{5}$$

where dist_{jn} denotes the geographical distance between j and n and t_{jns} is the sector-specific import tariff charged by n on imports from j. The binary variables lang_{jn}, and int_{jn} indicate whether j and n have an official language in common or are part of the same market, respectively. This last term is included in the specification of τ_{jns} since the price index for market n also includes firms from n itself. As a large body of research shows that border effects are quantitatively important, ignoring them would significantly underestimate the trade cost advantage of domestic firms (see McCallum, 1995; Anderson and van Wincoop, 2003).

With these assumptions, we obtain our empirical measure for the crowdedness of market n as

$$P_{ns}^{1-\sigma_s} = CR_{ns} = A_s \left[\sum_j ex_{jn}^{\alpha_1(\sigma_s-1)} \nu_{jns}^{\alpha_2} uc_{js}^{\alpha_3(1-\sigma_s)} \tau_{jns}^{1-\sigma_s} \right] \tag{6}$$

where $A_s = \Pi_{z=1}^{7}\mu_z \times \left(\frac{\sigma_s}{\sigma_s-1}\right)^{1-\sigma_s}$ summarizes constant terms, ex_{jn} denotes the bilateral exchange rate between countries j and n, ν_{jns} is our proxy for the number of exporters from j to n in sector s, uc_{js} are unit labor costs in country j, sector s, and τ_{jns} is defined as in (5).[6]

While (6) has been derived from a specific economic model we believe that its intuitive appeal is more general. For example, we can use CR_{ns} to ask what will happen to firm-level exports to market n if the number of competitors active there increases (ν_{jns} up), their unit costs decrease (uc_{js} down) or the trade barriers protecting the market are lowered (τ_{jns} down).

Expression (6) requires estimates for the parameters A_s and $\alpha_1(1-\sigma_s)$ to $\alpha_7(1-\sigma_s)$. These can

[6] We also experimented with linearizations of P_{ns} via Taylor-series expansions to obtain alternative empirical measures of market crowding (see Baier and Bergstrand, 2009, for such an approach in a different context). However, our results proved to be sensitive to the particular choice of center for the expansions. All specifications also suffered from severe multicollinearity between the elements of the linearized version of P_{ns} and the other regressors of our export demand specification (see equation (9) below).

be obtained from estimating gravity equations under the same assumptions which have been made so far. To see this, first note that the value of total exports from j to n in sector s is given by

$$R_{jns} = \int_{i_{jns}} p_{ijn}^{1-\sigma_s} P_{ns}^{\sigma_s - 1} E_{ns} di$$

We show in the technical appendix to this paper (Section 2) that under the same assumptions entering the derivation of (6), this can be written as:

$$R_{jns} = A_s ex_{jn}^{\alpha_1(\sigma_s - 1)} \nu_{jns}^{\alpha_2} uc_{js}^{\alpha_3(1-\sigma_s)} \tau_{jns}^{1-\sigma_s} P_{ns}^{\sigma_s - 1} E_{ns}$$

Using our functional form assumption for τ_{jns} from (5) and adding a time dimension, we derive the following gravity equation (in multiplicative form):

$$R_{jnst} = \beta_0 ex_{jnt}^{\beta_1} \nu_{jnst}^{\beta_2} uc_{jst}^{\beta_3} \times \left[\text{dist}_{jnt}^{\beta_4} (1 + t_{jnst})^{\beta_5} e^{\beta_6 \text{lang}_{jnt} + \beta_7 \text{int}_{jnt}} \right] \times d_{nst} \times \varepsilon_{jnst} \tag{7}$$

where ε_{jnst} is an error term and d_{nst} are destination-sector-time fixed effects, capturing the term $P_{nst}^{\sigma_s - 1} E_{nst}$ for which we do not have an empirical counterpart yet.

Note that while our proxy for the number of exporters (v_{jnst}) also contains information on trade flows, this information does not vary at the sectoral level. Indeed, this is the motivation for using aggregate rather than sectoral level trade shares when deriving v_{jnst} (see above). The use of domestic establishments as the second part of our exporter proxy further reduces problems related to an automatic correlation between R_{jnst} and v_{jnst}. We further explore this issue in the technical appendix to this paper (Section 3.1). As we show there, the results reported in Section 4 below are robust to different ways of proxying for the number of exporter and do not rely on the use of trade flow information in the derivation of v_{jnst}.

We estimate (7) by Poisson QMLE (see Wooldridge, 2002, chapter 19.2). As discussed by Santos-Silva and Tenreyro (2006), this estimation method yields consistent parameter estimates of log-linear models in the presence of heteroskedasticity (which, as these authors point out, is a key feature of trade data). It thus addresses an important source of inconsistency in existing estimation methods for gravity equations, such as OLS.

We use data on sectoral exports for all countries in our sample in 1992-2003.[7] Results are shown

[7] See Section 3 for details on our data. We pool data across four three-year periods in the regressions for comparability with the later firm-level regressions (see Section 3). To estimate (7), we need to convert trade flows into a common currency (U.S. dollars). Accordingly, the relevant exchange rate on the right-hand side is the exchange rate between exporter j's currency and the U.S. dollar. Note that it is the assumption of imperfect pass-through in our proxy for e_{jn}

in Table 1. Column 1 reports coefficient estimates from a regression pooling the data across industries and thus estimating a single coefficient for each of the required parameters. Columns 2-4 summarize estimates of sector-by-sector regressions by displaying the median, minimum and maximum coefficient estimates along with the corresponding t-statistics.

< Table 1 about here >

Overall, our results are very much in line with previous gravity equation estimates (see Disdier and Head, 2008). Distance has a significantly negative influence on bilateral trade while sharing a common language or being part of the same market all have a positive impact. Besides these more traditional determinants, the additional variables suggested by our model also have the expected sign and are highly statistically significant. In the pooled regression, a 1% increase in the exporter's unit labor costs reduces exports by around -0.29% while a 1% increase in the number of exporters is associated with 0.77% more exports.

We use our estimates from (7) to obtain the required parameter values in (6) as $A_s = \hat{\beta}_0$, $\alpha_1(1 - \sigma_s) = \hat{\beta}_1$, etc. Thus,

$$R_{nst} = \hat{\beta}_0 \left(\sum_j ex_{jnt}^{\hat{\beta}_1} \nu_{jnt}^{\hat{\beta}_2} uc_{jt}^{\hat{\beta}_3} \times \left[\text{dist}_{jn}^{\hat{\beta}_4} (1 + t_{jnt})^{\hat{\beta}_5} e^{\hat{\beta}_6 \text{lang}_{jn} + \hat{\beta}_7 \text{int}_{jn}} \right] \right) \tag{8}$$

For the main part of the analysis, we calculate (8) using the parameter estimates from the pooled regression (column 1 of Table 1). This is because these estimates have a much higher degree of precision than the sectoral-level estimates (which are often insignificant for a large fraction of industries). Section 4.2 presents results for robustness checks using the sectoral coefficient estimates.

Other variables Finding proxies for the remaining variables in (1) and (3) is straightforward. Total expenditure E_{ns} in market n, sector s, is proxied by total absorption, i.e., local production plus imports minus exports. For sector-specific trade costs between Italy and market n (τ_{in}), we use a similar assumption as before, i.e., $\tau_{in} = \mu_{I3}\text{dist}_{in}^{\gamma_3} \times \mu_{I4}(1 + t_{in})^{\gamma_4}$. We have dropped the indicator for common language since it is almost perfectly collinear to the regression's constant (the only other country which has Italian as an official language is Switzerland). Since we only consider exports, we further excluded the dummy for intranational trade.

To proxy firm-specific marginal costs c_i, we use two approaches. In analogy to our earlier assump-

which allows for a separate role for exchange rates in (7).

tions, we first consider the case $c_i = \mu_{I2}\left(\frac{w_i}{VA_i}\right)^{\gamma_2} = \mu_{I2}\,(uc_i)^{\gamma_2}$ where w_i denotes the total wagebill of firm i, VA_i is a firm's value added and uc_i its unit labor costs.[8] We will also estimate specifications with firm-by-year fixed effects (which capture c_i) to show that our results do not depend on this specific assumption.

Finally, we require an empirical counterpart for the initial setup costs F_{in} from equation (3). Our proxy for F_{in} should affect the export entry decision but not the value of exports. We use two different variables which arguably fulfil this property. The first is a firm's distance to Milan. Since Milan is Italy's business capital and learning about export markets happens in large part through contact with other exporting firms, proximity to Milan should lower F_{in}.[9] Secondly, we use an indicator for whether a firm is credit constraint or not. In most industries, setup costs have to be paid before any exports can take place and thus cannot be paid out of current export revenues. Since these initial investments can be considerable, credit is needed to finance them upfront (Roberts and Tybout, 1997).

2.3 Empirical Specifications

With these empirical proxies, we arrive at our baseline estimation equation

$$r_{int} = \begin{cases} \gamma_0 ex_{int}^{\gamma_1} uc_{it}^{\gamma_2} \text{dist}_{in}^{\gamma_3} (1+t_{int})^{\gamma_4} E_{nst}^{\gamma_5} CR_{nst}^{\gamma_6} \eta_{1inst} \text{ if } D_{int} \geq 0 \\ 0 \text{ otherwise} \end{cases} \tag{9}$$

where $D_{int} = \delta_0 ex_{int}^{\delta_1} uc_{it}^{\delta_2} \text{dist}_{in}^{\delta_3} (1+t_{int})^{\delta_4} E_{nst}^{\delta_5} CR_{nst}^{\delta_6} F_{int}^{\delta_7} \eta_{2inst}$, and η_{1inst} and η_{2inst} are error terms.

3 Data and Descriptive Statistics

Firm-level data on exports and other firm characteristics come from a survey conducted every three years by Capitalia on a representative sample of Italian manufacturing firms.[10] We use the four most recent waves of the survey which contain information for four three-year periods between 1992-2003 (1992-1994, 1995-1997, 1998-2000, 2001-2003). The main variables we use are the value of exports

[8] A sufficient condition for this approximation to hold exactly is that the short-run value added production function (i.e., after set-up costs have been incurred) is Cobb-Douglas with constant returns to scale. We also need the cost of capital to be either identical across firms or proportional to wages or total factor productivity. Note that most of our regressions will include industry-by-year fixed effects so that the cost of capital can vary across industries.

[9] We acknowledge that this variable might also capture the distance between firms and their export markets more generally, and thus influence the value of exports to these markets. This is because our distance variable measures average (not firm-specific) distance between Italy and its export markets (see Section 3).

[10] Data from the Capitalia survey have been extensively used by other authors to study different aspects of the Italian economy (e.g., Benfratello et al., 2010; Caggese and Cuñat, 2008; Hall et al., 2009). Most closely related to the present work are a number of papers concerned with the internationalization strategies of Italian firms and the corresponding effects on variables such as productivity (e.g., Castellani, 2002; Casaburi et al., 2007; Vannoni and Razzolini, 2008).

by destination and unit labor costs (a firm's wage bill divided by value added). Export destinations are grouped by eight main geographical areas, plus Italy itself.[11] After dropping observations with missing firm or export information, we obtain a sample of 3,628 firms, 8,087 firm-year pairs and 64,256 firm-year-destination trade flows.

The country-level data required for the calculation of our market crowding measure come from a number of sources. Sector-level information on value added, the total wage bill and the number of establishments are from UNIDO's Industrial Statistics Database and the OECD's Structural Analysis Database, completed with national sources in the case of missing data. Bilateral exchange rates are from the IMF's International Financial Statistics. The trade data we use for estimating the model's parameters are provided by CEPII (2005). Data on bilateral tariffs, distances, and common official languages is also from CEPII (2005, 2006). After dropping countries with missing data we obtain a sample of 75 countries which on average accounted for more than 90% of world GDP and trade in our sample period 1992-2003

We calculate our measure of market crowding using absorption-weighted averages for the bilateral variables in (8). For example, the distance between the United Kingdom (j) and NAFTA (n) is $\text{dist}_{jn} = \sum_{m \epsilon n} \text{dist}_{jm} \times \text{share}_{mn}$, where share_{mn} is the share of country m in total absorption of NAFTA and $n = \{\text{USA, Canada, Mexico}\}$. We use the same approach for obtaining bilateral distances and tariffs for Italian firm-level exports in (9).[12]

The technical appendix to this paper contains further details about the above data and descriptive statistics on our firm-level variables and the market crowding measure, CR_{nst}. It also explains the construction of auxiliary variables such as our proxies for F_{in} (distance to Milan, dummy for credit constraints).

4 Econometric Results

4.1 Baseline Specification

We start by estimating a Heckman selection model based on a log-linearized version of equation (9). We choose Heckman as our baseline since a selection equation approach follows immediately from our theoretical model. Below, we will check the robustness of our results to alternative estimation

[11] These are Europe (EU15 excluding Italy), other European countries (including Russia and Turkey), NAFTA (United States, Canada and Mexico), Central and South American countries, China, other Asian countries (excluding China), Africa, and Australia and Oceania.

[12] We use total rather than sectoral absorption and compute shares for 1989-1991 to reduce endogeneity problems.

techniques such as Poisson QMLE.

Column 1 of Table 2 presents results without exclusion restrictions (i.e., assuming that $\delta_7 = 0$). Identification thus relies on the nonlinearity of the inverse Mills ratio (Wooldridge, 2002). Column 2 uses as the exclusion restriction our credit constraints indicator, whereas column 3 proxies the setup costs of exporting (F_{in}) by the distance of the exporter to Milan.

As shown in Table 2, the results are not very sensitive to the choice of exclusion restriction. In all cases, foreign absorption enters significantly with a positive contribution while market crowding, distance to the export market and firm-level unit labor costs show the expected negative sign. Looking at the selection equations, a similar pattern holds for the decision to export to a specific market. Higher absorption and lower unit labor costs raise the probability that a firm is active in market n, while distance and market crowding reduce it. Note that the excluded variables (distance to Milan, dummy for credit constraints) are also significant and have the expected sign – both lowering the probability of export market entry.

To get an impression of the overall impact of the regressors, we also report marginal effects evaluated at the sample mean. As seen, the combined effect of a 1% increase in the level of crowdedness of a foreign market is a -0.21% to -0.22% decrease in firm-level exports there. For foreign absorption, a 1% increase leads to an increase in exports of 0.50% to 0.51%. Marginal effects for a 1% increase in the other variables are +0.65% to +0.66% for exchange rates, -3.19% to -3.38% for tariffs, -0.74% to -0.75% for bilateral distance and -0.31% to -0.36% for unit labor costs.

< Table 2 about here >

4.2 Robustness Checks

Table 3 reports a number of robustness checks on our initial results.[13] In column 1, we exclude firms which do not export to any foreign market in a given period t. We note that our approach quite naturally allows for non-exporters – these are firms with unit costs which are too high to profitably enter any foreign market. Nevertheless, the decision to export at all might be fundamentally different from the decision to export to any given market. As our results show, excluding non-exporters slightly changes some of the coefficient estimates. The qualitative nature of our baseline results remains very

[13] Since Table 2 suggests that results are robust to the absence of exclusion restrictions, we report results for Heckit estimations without such restrictions. This maximizes the number of available observations and increases comparability with the alternative estimation techniques reported below. To save space, we also only report marginal effects from now. Full results are available from the authors upon request.

much intact, however, with the coefficient on CR being almost identical to before.

Columns 2 and 3 control for conditions on other markets. While we ruled out such an influence in the earlier theoretical part by assuming segmented markets and constant marginal costs, third-market conditions might of course be relevant in the data.

Column 2 includes an absorption-weighted average of the crowdedness of all eight foreign markets,

$$CR_{RoW,st} = \sum_n \text{share}_{ns} \times CR_{nst}$$

where share_{ns} is the average share of market n in the overall absorption of industry s over the period 1992-2003. Column 3 includes market crowding in Italy itself, calculated in the same way as CR_{nst} for all other markets. In both specifications, we also add total industry absorption in Italy and the rest of the world, respectively, as an additional control.

As the results show, the sign of these variables is as expected. Higher demand in Italy or the rest of the world reduces exports to any given market whereas higher levels of market crowding in Italy itself or the rest of the world increase exports. Thus, firms do indeed seem to take conditions on third markets into account in their export decisions. In both specifications, however, foreign market crowding remains statistically and economically significant, demonstrating that our results are qualitatively robust to the inclusion of third-market controls.

< **Table 3 about here** >

Columns 4 and 5 report results with different sets of fixed effects to further address concerns about omitted variable bias. Column 4 adds destination-year fixed effects and column 5 uses industry-year specific effects. Using destination-year fixed effects reduces the magnitude of the destination-varying regressors while using industry-year fixed effect has the opposite impact. A possible explanation is that the cross-destination variation in our regressors is more important than the cross-industry variation and that measurement error thus tends to bias results more strongly towards zero in the destination-year fixed effects specification. In both cases, however, all regressors retain their sign and significance, indicating that the earlier results are not relying on a single dimension of the data. Note that the use of industry-year fixed effects also controls for the influence of alternative export markets which were found to be important above. We use this set of fixed effects for most of our remaining empirical results and the counterfactual experiments in Section 5.[14]

[14] Using industry-by-year fixed effects implies that identification relies on cross-destination variation in the data. In

In column 6, we recalculate our measure of market crowding using the sector-specific estimates from Table 1. While estimation precision is much lower, they are closer to the theoretical model from Section 2. This is because elasticities of substitution σ_s are likely to vary across sectors which in turn will influence the degree of market crowding. As shown, the qualitative picture of the previous regressions stays intact when allowing for this additional variation.

Columns 7-11 report results for alternative estimation techniques. Columns 7-10 show results obtained via the QML Poisson estimator discussed in Santos-Silva and Tenreyro (2006) and Wooldridge (2002, chapter 19.2). This estimation method does not allow to disentangle the extensive and intensive margin of export decisions suggested by the model. However, it has the advantage over the previous Heckit estimates of not imposing distributional assumptions on the error structure – a correct specification of the conditional mean is sufficient to obtain consistent estimates. Poisson further allows estimation of results with firm-by-year fixed effects since (unlike Heckit) it is not susceptible to incidental parameter problems (see Wooldridge, 2002).

For comparison with the earlier Heckit results, we estimate Poisson models with identical sets of fixed effects (year, destination-year, and industry-year). As shown in columns 7-9, the Poisson QMLE coefficient estimates are generally larger in magnitude than the corresponding Heckit marginal effects. Overall, however, the qualitative picture of the earlier results remains very much intact. Further adding firm-by-year fixed effects (column 10) leaves the coefficient on CR and the other regressors almost unchanged.[15]

Finally, column 11 shows estimates from a Poisson instrumental variables (IV) regression (see Mullahy, 1997). Using Poisson IV allows us to further address concerns about omitted variable bias. It also addresses bias arising from measurement error which is likely to be relevant in our context, given that we had to use a number of proxy variables when constructing the empirical counterpart to the CES price index, CR.

Similar to the linear IV estimator, IV Poisson requires an instrument that is correlated with CR but not with potentially omitted variables, or the measurement error part of CR (Mullahy, 1997). Here, we use bilateral distances and the common language indicators from (5) as instruments. Since they are

our view, this is also closer to the theoretical framework from section 2 which models firm-export decisions to different destinations within a given industry. All qualitative results in the remainder of the paper carry through under alternative sets of fixed effects (year or destination-year).

[15] As we discuss in the data section of the technical appendix to this paper, the time dimension of our panel is very short (two adjacent three-year periods for most firms). It is thus not possible to include firm-destination specific fixed effects since the low time-series variation would substantially lower the signal-to-noise ratio in our data, aggrevating measurement error bias. However, as long as there is no significant bias arising from omitted firm-market specific factors, our coefficient estimates will still be consistent even when relying on other fixed effects combinations. We can thus use these estimates for the purpose of counterfactual experiments as we will do below.

both part of CR, they will by construction be correlated with our crowdedness measure. They are also likely to be valid instruments, given that they are clearly exogeneous to the export decision of Italian firms and probably measured with none or very little error (in contrast to other elements of CR, such as the number of exporters or their marginal costs, n_{jns} and c_i). Specifically, we instrument CR with the weighted average distance of a market n from all other countries j, and with the weighted average of common language ties with all countries exporting to n. That is, $\text{avgdist}_n = \sum_j \text{share}_j \times \text{dist}_{jn}$ and $\text{avglang}_n = \sum_j \text{share}_j \times \text{lang}_{jn}$, where share_j is the share of exporter j in world absorption, calculated in the pre-sample period 1989-1991 to preserve exogeneity of the instruments.

The results in column 11 suggest that biases arising from omitted variables and measurement error are not a major concern for our results. Indeed, the IV results are very close to the corresponding Poisson estimates from column 7. We now estimate an elasticity of exports with respect to market crowding of -0.47 compared to -0.54 in our baseline Poisson specification.[16]

4.3 Alternative Measures of Export Market Competition

We also present results for three non-structural measures of export market crowding. First, we use the average trade-weighted import tariff of market n:

$$\text{AvgTar}_{nst} = \sum_j (\text{share}_{jns} \times \text{tariff}_{jnst}) \tag{10}$$

where tariff_{jnst} is the average tariff imposed in market n, sector s, on imports from country j. These tariffs are weighted by the average sector-specific import share of country j in market n over the entire period 1992-2003 (share_{jns}).

Second, we construct a measure based on the Herfindahl index for market n, sector s. We do not have data on the market shares of individual firms. Instead, we assume that total exports from j to n are equally split among exporters in j. Thus,

$$\text{Herf}_{nst} = \sum_j n^H_{jnst} \left(\frac{\text{mshare}_{jnst}}{n^H_{jnst}} \right)^2$$

where mshare_{jnst} is the share of country j in total absorption of market n, sector s, period t. We use a

[16] We note that this change in coefficient magnitudes is likely to capture a number of factors. While the presence of random measurement error would imply higher IV estimates, non-random measurement error and omitted variable bias might well bias the original coefficient estimates away from zero. The key point is that the two coefficient estimates are close and the overall magnitude of the bias is thus unlikely to be very large. In unreported results, we also experimented with other fixed effects combinations. Again, results for IV Poisson estimates were of similar magnitude to the corresponding Poisson regressions.

similar proxy for the number of exporters as the one described in Section 2, $n^H_{jnst} = \text{est}_{js} \times \text{share}_{jns}$.[17] If exports from j to n are equally distributed among exporters, each exporter will have a market share of $\text{share}_{jnst}/n^H_{jnst}$. Squaring this share, multiplying by n^H_{jnst} and summing over all countries exporting to n (including domestic firms) then yields the Herfindahl index for the respective market and industry.

Third, we simply count the number of firms active in a given market, again using $n_{jnst} = \text{est}_{js} \times \text{share}_{jns}$ as a proxy. This captures the idea that a higher number of firms active in a location implies a higher degree of competition, ceteris paribus:

$$N_{nst} = \sum_j n^H_{jnst}$$

Table 4 present the results for these three alternative measures. The tariff variable AvgTar_{nst} is significant and has the expected sign in column 1. Ceteris paribus, higher average destination market tariffs should increase exports since foreign competitors will find access to that market more difficult (controlling for the tariffs faced by Italian exporters themselves). However, this result is not robust to the inclusion of industry fixed effects in column 2 where the coefficient on AvgTar is statistically indistinguishable from zero.[18]

< Table 4 about here >

A similar pattern holds for the Herfindahl index (Herf_{nst}). It has the expected positive sign in column 3, indicating that more concentrated and thus presumably less competitive markets attract more Italian exports, ceteris paribus. Note, however, that this effect is not statistically significant. Controlling for industry-year fixed effects yields a coefficient estimate which is statistically significant at the 10%-level, but which has the wrong (negative) sign.

Finally, the count of active firms, N_{nst}, enters with a positive sign in both specifications, although the corresponding coefficient estimate is statistically indistinguishable from zero in the year-fixed-effects-only regression (see columns 5-6). These results indicate that markets with more competitors actually attract *more* Italian exports, ceteris paribus, confirming the difficulties of non-structural

[17] This corresponds to our proxy $n_{jns} = \mu_2 \left(est_{js} \times share_{jn}\right)^{\alpha_2}$ from section 2 with parameters $\alpha_2 = \mu_2 = 1$. Note that we can now use sectoral level variation in the trade share variable ($share_{jns}$) since we no longer need to be concerned with an automatic correlation with the dependent variable in our gravity regressions (compare the discussion in section 2). Using $share_{jn}$ as in the original proxy does not affect the conclusions below.

[18] We note that these results do not contradict previous findings in the literature on the role of access to foreign markets for firm-level decisions, which usually suggest a strong role for foreign tariffs (e.g., Lileeva and Trefler, 2009). Indeed, we also find that foreign tariffs are a significant determinant of Italian exports (see Tables 2-4). Compared to this direct effect, the indirect effect of foreign tariffs working through better access for other firms is likely to be of a second order magnitude only, as our results confirm.

measures to yield robust and plausible predictions with regards to the impact of market crowding on firm-level exports. One way to interpret these findings is that more information about the firms competing with Italian exporters in a given market needs to be added to obtain a robust indicator for the crowdedness of a market. For example, the measure we proposed in Section 2 combines information on the number of competitors, their productive efficiency and the access barriers they face in a given market.

4.4 Firm Heterogeneity

There are several apriori reasons why one might expect the effect of market crowding to vary across firms. This section investigates this issue further. First, one might expect vertically differentiated firms to be less affected by the degree of crowdedness of a foreign market. Investing in higher product quality is one important way for Italian firms to increase vertical differentiation and thus to reduce the impact of competition from other exporters and local firms.

Secondly, firms are likely to be less affected by foreign market conditions if they belong to international networks within multinational entreprises. For example, they might sell goods abroad using different distribution channels or sell to other firms within the same group. This will make them less susceptible to the influence of foreign market crowding.

A final source of potential heterogeneity comes from the type of ownership. In Italy, a substantial share of firms (of any size) are owned and managed by families. Barba Navaretti, Faini and Tucci (2008) show that such firms tend to export less and to less distant markets than publicly owned firms.

Table 5 presents results for regressions allowing for heterogeneity by interacting proxies for the above firm characteristics with foreign market crowding. We classify a firm as being vertically differentiated and/or producing higher quality goods if it engages in R&D activity (i.e., employs workers in R&D). "Multinational companies" are firms that are either foreign owned or have affiliates abroad. "Family firms" are firms that are managed by the owner or a member of its family.[19]

< Table 5 about here >

Columns 1-3 introduce these characteristics one by one. In all cases, the interaction term between firm characteristics and market crowding enters with a positive coefficient. That is, firms engaging in R&D, firms which are not family-owned and those which are part of an MNE tend to be less affected

[19] The technical appendix to this paper contains summary statistics on these variables as well as further details about their construction. Note that we code the family-firm dummy as 0 for family firms and 1 for other firms.

by the crowdeness of foreign markets. However, the effects are relatively small in magnitude and only statistically significant for R&D and multinational status.

In Column 4, we include all three characteristics and their interactions with the CR variable in the same specification. As before, the marginal effects of the interaction terms are all positive albeit only statistically significant for R&D. Note that the firm characteristics which we interact with CR are not mutually exclusive. Our results thus indicate that the firms most affected by foreign market crowding are family firms that are not part of multinationals and do not employ R&D workers. On the other end of the range are multinationals not managed by a family and employing workers in R&D. According to our results, the overall average impact of foreign market crowding is around 0.07 log points less important for this latter group than for the former. This difference is statistically significant although economically not very large – market crowding clearly matters for all types of firms. In the following counterfactual experiments we thus focus on our baseline results from sections 4.1 and 4.2.

5 Quantitative Importance of Results

We now turn to an evaluation of the quantitative importance of our results. We do so by performing a series of counterfactual experiments. That is, we set the elements of the right-hand side of (9) to new counterfactual values, compute the predicted exports and compare them to their original value.

Specifically, let $r_{int} = E(r_{int}|X_{int})\varepsilon_{int}$ denote the original value of exports and $\hat{r}_{int} = E(r_{int}|\hat{X}_{int})\varepsilon_{int}$ the exports under the counterfactual values of the regressors X. The percentage change in exports across all firms between the counterfactual and the actual scenario is then

$$\frac{\hat{R}_{it} - R_{it}}{R_{it}} = \sum_{i=1}^{I}\sum_{n=1}^{N} s_{it}s_{int}\frac{E(r_{int}|\hat{X}_{int}) - E(r_{int}|X_{int})}{E(r_{int}|X_{int})}$$

where s_{it} and s_{int} are, respectively, the share of firm i in total actual exports (R_{it}) and market n's share in firm i's actual exports (r_{it}). Note that we require values for $E(r_{int}|.)$ rather than for $E(\ln r_{int}|.)$ for these calculations. For our Heckit estimates, the expected value of r_{int} in levels is given by (see Dow and Norton, 2003):

$$E(r_{int}|X) = \Phi\left(X_1\hat{\gamma}_{sel} + \hat{\rho}\hat{\sigma}_1\right)\exp\left(X_2\hat{\gamma}_{out} + 0.5\hat{\sigma}_1^2\right)$$

where $\hat{\sigma}_1$ is the estimated variance of the outcome equation's error term and $\hat{\rho}$ the estimated coefficient of correlation between outcome and selection equation residuals. X_1 and X_2 denote the variables in the selection and outcome equation, respectively, and $\hat{\gamma}_{sel}$ and $\hat{\gamma}_{out}$ are the corresponding coefficient estimates. Note that the Poisson regressions directly give us $E(r_{int}|X)$ so that no transformation is necessary.

We start our counterfactuals by setting the regressors $X = \{ex, uc, t, E, CR\}$ on the right-hand side of (9) to their values lagged by one period. We do so separately for each of the regressors. This allows us to calculate the growth rate of total exports in the absence of, for example, demand growth ($\hat{E}_{nst} = E_{nst-1}$), changes in exchange rates ($e\hat{x}_{nt} = ex_{nt-1}$), or unit labor costs ($u\hat{c}_{it} = uc_{it-1}$). Table 6 shows results for all regressors except distance which is time-invariant. We report geometric averages of growth rates across periods, expressed in %-changes per year. These figures thus tell us by how much more or less Italian exports would have grown per year in the absence of any changes in, say, absorption or unit labor cost over the sample period 1992-2003.[20]

We note that these counterfactuals are not general equilibrium in nature. For example, lower demand growth is likely to result in a reduction in the number of foreign competitors active on Italian export markets. It might also result in lower world-wide demand for manufacturing inputs and thus lower unit labor costs of Italian producers.

With these caveats in mind, we turn to a discussion of our results. The Heckit estimates allow us to analyze the effect of the above counterfactual changes on the probability of selection into exporters status (the "extensive margin") and the value of exports taking the probability of selection as given (the "intensive margin"). We thus report three counterfactual growth rates of exports. First, we only use counterfactual values of the regressors X_1 in the outcome equation (the intensive margin, column 1). Next, we only set the regressors in the selection equation, X_2, to their new values (the extensive margin, column 2). Finally, we change both X_1 and X_2 which gives us the total effect of the counterfactual change.

< **Table 6 about here** >

According to the counterfactuals based on our Heckit estimates, absorption is by far the most important determinant of export growth. Keeping absorption constant would have reduced exports by around -4% per year or by 42% over the sample period. Not unexpectedly, the biggest contribution to this overall figure comes from the EU15 (excluding Italy). Holding absorption growth there constant would have meant -1.3% less exports per year. The second and third most important markets with

[20] The reason for taking this approach is that we do not observe all firms in all periods or even in the first and last period. Otherwise, we could have simply set regressors in 2001-2003 to their 1992-1994 values and computed counterfactual yearly growth rates.

regards to demand growth were other European countries (-0.9%/year) and NAFTA (-0.2%/year).

Exchange rate variations and changes in unit labor costs were also important. Holding unit labor costs fixed would have allowed Italian exports to grow by around 1.4%/year more rapidly. Keeping exchange rates unchanged would have increased Italian exports by 2% per year. This relatively large figure seems to be mainly due to the large scale devaluations of large South American importers (Brazil, Argentina) over the sample period. This is evident from the next two lines where we disaggregate results by allowing South and Central American exchange rates to vary but holding all other exchange rate fixed – as well as the other way around.

Turning to the remaining regressors, the roles of tariffs and market crowding are less significant. In the absence of any further tariff reductions after 1992, Italian export growth would have been -0.3% per year lower. The impact of freezing the level of market crowding is actually the smallest among all regressors. Holding it constant would have increased exports by only around 0.2% per year or around 1.8% over the entire sample period.

The same qualitative picture reappears when looking at the results based on the Poisson coefficient estimates. These estimates imply a somewhat stronger impact of market crowding (0.34% per year or 3.1% over the sample period) but most of the other factors also become more important. For example, an absence of demand growth now would have reduced exports by -5% per year or around 55% over the sample period. Thus, changes in market crowding were an order of magnitude less important than changes in foreign demand conditions.[21]

Of course, these aggregate figures might hide substantial variation across the components of CR which could cancel each other out. We thus also report the impact of the various components of CR – number of exporters, foreign tariff, exchange rates and unit labor costs. That is, for each one of these components we recompute our market crowding measure while only holding this particular variable constant over time. As the results indicate, the impact of changes in the number of exporters, foreign tariffs and unit labor costs all worked in the same direction – each contributed towards a (small) reduction in Italian exports.

Another possible decomposition of the impact of market crowding is to look at the role of individual countries. We do so by returning to one of the motivating questions for this paper and ask whether Italian exports would have grown faster or more slowly in the absence of China's integration into the world economy. To this end, we first fix the contribution of China to our CR measure. That is, we

[21] Note that the coefficient estimates for CR used in this section cover most of the range of estimates in Tables 2 and 3 (from -0.48 for Heckit up to -0.67 for Poisson). Results using any of the other sets of coefficient estimates from these tables are qualitatively similar and available upon request.

hold constant the number of Chinese exporters, their unit labor costs, and the exchange rates and tariffs they face. Since CR is a sum over all countries in our sample, China's impact could in principle be bigger than the aggregate figure of -0.2% per year presented above. Secondly, we do the same with China's absorption growth and its external tariffs and exchange rate facing Italy. As the results in Table 7 indicate, the role of China is not very important in our sample period and if anything positive. China's integration into the world economy meant more competition for Italian exporters but this effect is negligible (0.02% per year).[22] Furthermore, it is dominated by increased exporting opportunities to the large Chinese market. The absence of absorption growth in China would have lowered Italian exports by -0.13% per year and freezing Chinese import tariffs at their 1992 level would have contributed another -0.09% per year. Overall, we estimate that in the absence of changes in China and its integration into the world trading system, Italian exports would have grown by -0.2% per year less quickly. Again, results using the Poisson estimates are very similar.

< Table 7 about here >

As a final counterfactual, we ask what Italian exports would have been if absorption growth in the EU15 had been as rapid as in the rest of the world – i.e., on average 1.5% per year higher than it has been. As the last row in Table 7 shows, the slow growth in demand in Italy's main market is an order of magnitude more important than the emergence of China. Bringing EU15 demand growth up to the world average would have increased Italian exports by up to 0.8% per year.

6 Conclusions

This paper examined the role of foreign market conditions for firm-level exports. Given the growing share of exports in manufacturing production it is of key interest for both academic and economic policy debates to obtain a better understanding of how levels of demand and competition intensity in foreign markets affect export performance.

We started by constructing a simple firm-level gravity model to derive an econometric specification linking destination-specific exports to firm characteristics, foreign demand and the degree of competitiveness or "crowdedness" of foreign markets. This latter variable is a measure of the number and efficiency of firms competing in a given market and the barriers impeding their access, such as tariffs

[22] This result echoes Hanson and Robertson's (2008) finding that even those developing countries specialized in manufacturing goods were only very modestly affected by China's export expansion.

or physical distance.

We estimated this specification on a large sample of Italian manufacturing firms for the period 1992-2003. Having shown that market crowding has a robust negative impact on firm-level exports across a wide range of specifications, we used our estimates to evaluate the quantitative importance of market crowding.

Our main specification indicates that increased numbers and efficiency of foreign firms combined with a better overall accessibility of destination markets have reduced Italian exports by around 0.2% per year or 1.8% over the sample period. This is similar to the effects of tariff reductions for Italian firms (+0.3% per year) but smaller than the impact of higher unit labor costs (-1.4% per year) and less favorable exchange rates (-2.0% per year). By far the most important determinant of export performance was foreign demand growth, raising Italian exports by up to 5% per year or 55% over the sample period.

Our results also indicate that the role of China in explaining Italian export performance is small and if anything positive. Stronger competition from China marginally lowered Italian exports but this was overcompensated by Chinese demand growth and tariff reductions, yielding an overall positive effect on export growth of 0.2% per year. Much more important was the fact that demand on Italy's main export market, the EU15, has grown more slowly during 1992-2003 than in the rest of the world. Bringing demand growth in the EU15 up to the world average would have increased Italian exports by up to 0.8% per year.

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Table 1: Estimation of Parameters – Gravity Equation

	Dep. Var. Bilateral Exports			
Specification	Pooled	Sector-specific		
		Min	Median	Max
ln(exchange rate)	0.031	-0.075	0.028	0.127
	(3.58)**	(2.10)*	(0.63)	(3.61)**
ln(exporters)	0.772	0.346	0.776	0.991
	(53.86)**	(4.41)**	(11.15)**	(19.23)**
ln(unitcost_exporters)	-0.290	-1.134	-0.383	0.130
	(8.23)**	(5.46)**	(2.11)*	(0.50)
ln(distance)	-0.250	-0.896	-0.309	-0.044
	(7.55)**	(4.47)**	(1.80)+	(0.17)
ln(1+tariff)	-1.906	-11.449	-0.495	1.127
	(2.81)**	(4.52)**	(0.24)	(0.93)
Common language	1.009	0.225	1.024	1.759
	(10.16)**	(0.53)	(2.46)*	(5.02)**
Internal trade flow dummy	0.281	0.064	0.438	1.909
	(4.92)**	(0.20)	(1.73)+	(7.30)**
Fixed Effects	Importer-Industry-Year	Importer-Year	Importer-Year	Importer-Year
Observations	73476	2538	2736	2772

Notes: Table displays coefficients for Poisson QMLE (t-statistics in brackets, based on standard errors clustered on exporter-importer-industry pairs in column 1 and exporter-importer pairs in columns 2-4). Column 1 pools all sectors while columns 2-4 present results for sector-specific regressions. For each regressor, we display the minimum, median and maximum coefficient estimate across regressions, as well as the minimum, median and maximum number of observations (estimates and number of observations in a given column can thus come from different regressions). +, * and ** signify statistical significance at the 10%, 5% and 1% levels, respectively.

Table 2: Baseline Results - Heckit

	(1)			(2)			(3)		
	ln(exp)	d(exp>0)	ME	ln(exp)	d(exp>0)	ME	ln(exp)	d(exp>0)	ME
ln(CR)	-0.235	-0.095	-0.221	-0.177	-0.096	-0.209	-0.223	-0.089	-0.207
	(4.86)**	(4.46)**	(5.06)**	(3.80)**	(4.39)**	(4.73)**	(4.66)**	(4.12)**	(4.72)**
ln(ex. rate)	0.651	0.292	0.662	0.59	0.294	0.653	0.635	0.287	0.652
	(14.58)**	(12.98)**	(14.05)**	(12.89)**	(12.83)**	(13.83)**	(14.02)**	(12.64)**	(13.79)**
ln(absorption)	0.506	0.224	0.510	0.449	0.225	0.499	0.491	0.219	0.498
	(14.99)**	(14.89)**	(16.40)**	(12.76)**	(14.57)**	(15.88)**	(14.29)**	(14.38)**	(15.95)**
ln(1+tariff)	-2.455	-1.608	-3.379	-2.057	-1.621	-3.318	-2.254	-1.523	-3.191
	(5.79)**	(7.80)**	(8.09)**	(4.96)**	(7.56)**	(7.80)**	(5.43)**	(7.37)**	(7.66)**
ln(distance)	-0.577	-0.352	-0.749	-0.508	-0.354	-0.738	-0.562	-0.355	-0.753
	(15.34)**	(21.41)**	(21.43)**	(12.22)**	(21.19)**	(21.22)**	(14.10)**	(21.47)**	(21.50)**
ln(unitcost)	-0.765	-0.089	-0.338	-0.749	-0.076	-0.310	-0.79	-0.099	-0.361
	(13.85)**	(4.01)**	(7.43)**	(14.08)**	(3.32)**	(6.71)**	(13.97)**	(4.41)**	(7.86)**
Credit constraint dummy					-0.092	-0.159			
					(4.84)**	(4.79)**			
Travel time to Milan								-0.033	-0.058
								(8.73)**	(8.47)**
Fixed effects		Year			Year			Year	
Observations		64256			61592			62312	

Notes: Table displays coefficients for Heckman selection models (t-statistics in brackets, based on standard errors clustered on industry-destination-years). For each model, the table display results for outcome and selection equation, and marginal effects evaluated at sample means. ME denotes “marginal effect”. * and ** signify statistical significance at the 5% and 1% levels.

Table 3: Robustness Checks

	(1) ME	(2) ME	(3) ME	(4) ME	(5) ME	(6) ME	(7) ME	(8) ME	(9) ME	(10) ME	(11) ME
ln(CR)	-0.230	-0.305	-0.658	-0.175	-0.476	-0.275	-0.537	-0.538	-0.673	-0.703	-0.468
	(4.10)**	(4.58)**	(8.72)**	(3.71)**	(8.61)**	(6.50)**	(6.22)**	(6.61)**	(3.87)**	(4.56)**	(2.08)*
ln(ex. rate)	0.948	0.756	0.939		0.828	0.725	0.909		0.989	1.013	0.854
	(15.45)**	(14.36)**	(15.25)**		(18.73)**	(18.70)**	(9.36)**		(9.90)**	(11.31)**	(6.09)**
ln(absorption)	0.674	0.598	0.783	0.314	0.678	0.577	0.778	0.749	0.888	0.912	0.709
	(16.71)**	(15.18)**	(15.87)**	(7.67)**	(18.92)**	(18.36)**	(9.09)**	(6.32)**	(8.77)**	(10.34)**	(5.88)**
ln(1+tariff)	-4.778	-3.151	-3.792	-1.468	-4.022	-3.989	-2.506	-3.03	-2.496	-2.428	-2.781
	(9.33)**	(8.46)**	(9.27)**	(3.47)**	(13.17)**	(11.73)**	(2.66)**	(3.31)**	(4.03)**	(4.18)**	(4.65)**
ln(distance)	-1.155	-0.759	-0.817		-0.781	-0.742	-0.778		-0.817	-0.827	-0.659
	(25.95)**	(21.78)**	(22.98)**		(34.59)**	(34.85)**	(12.33)**		(13.91)**	(14.98)**	(13.84)**
ln(unitcost)	-0.325	-0.364	-0.377	-0.338	-0.656	-0.657	-0.396	-0.435	-0.806		-0.550
	(5.53)**	(8.22)**	(8.32)**	(7.43)**	(16.10)**	(16.11)**	(4.26)**	(4.93)**	(7.05)**		(3.38)**
ln(absorb_RoW)		-0.232									
		(7.05)**									
ln(CR_RoW)		0.062									
		(0.92)									
ln(absorb_Italy)			-0.411								
			(6.47)**								
ln(CR_Italy)			0.497								
			(7.82)**								
Fixed effects	Year	Year	Year	Destin.-Year	Industry-Year	Industry-Year	Year	Destin.-Year	Industry-Year	Firm-Year	Year
Estimation Method	Heckit	Heckit	Heckit	Heckit	Heckit	Heckit	Poisson	Poisson	Poisson	Poisson	Poisson IV
Observations	46400	64256	64256	64256	64256	64256	64256	64256	64256	64256	64256

Notes: Table displays coefficients and t-statistics based on standard errors clustered on industry-destination years. Estimation methods are Heckit (columns 1-6), Poisson (columns 7-10) and Poisson IV (column 11). The dependent variable is firm-destination exports. ME denotes "marginal effect". For Heckit, we report marginal effects evaluated at sample means. * and ** signify statistical significance at the 5% and 1% levels.

Table 4: Non-Structural Measures of Market Crowding

	(1) ME	(2) ME	(3) ME	(4) ME	(5) ME	(6) ME
ln(AvgTariff)	0.162	-0.007				
	(4.53)**	(0.30)				
ln(Herfindahl)			0.027	-0.105		
			(0.71)	(1.94)+		
ln(N_{nst})					0.002	0.165
					(0.05)	(3.08)**
ln(exchange rate)	0.672	0.549	0.558	0.479	0.537	0.471
	(12.00)**	(14.78)**	(11.23)**	(9.74)**	(11.49)**	(12.18)**
ln(absorption)	0.444	0.405	0.412	0.319	0.388	0.302
	(15.18)**	(21.04)**	(11.35)**	(6.65)**	(12.52)**	(8.16)**
ln(1+tariff)	-4.197	-3.431	-3.177	-3.420	-3.193	-3.228
	(8.41)**	(9.96)**	(7.17)**	(9.91)**	(7.10)**	(9.14)**
ln(distance)	-0.699	-0.700	-0.713	-0.680	-0.707	-0.626
	(18.49)**	(33.14)**	(19.88)**	(29.05)**	(19.23)**	(20.46)**
ln(unitcost)	-0.387	-0.660	-0.373	-0.665	-0.381	-0.664
	(8.60)**	(16.14)**	(8.11)**	(16.38)**	(8.36)**	(16.38)**
Fixed effects	Year	Industry-Year	Year	Industry-Year	Year	Industry-Year
Estimation Method	Heckit	Heckit	Heckit	Heckit	Heckit	Heckit
Observations	64256	64256	64256	64256	64256	64256

Notes: Table displays coefficients and t-statistics for marginal effects obtained via Heckit (t-statistics based on standard errors clustered on industry-destination years). The dependent variable is firm-destination exports. ME denotes "marginal effect". Marginal effects are evaluated at sample means. +, * and ** signify statistical significance at the 10%, 5% and 1% levels.

Table 5: Firm Heterogeneity

	(1) ME	(2) ME	(3) ME	(4) ME
Ln(CR)	-0.513	-0.478	-0.488	-0.518
	(9.14)**	(8.48)**	(8.80)**	(9.15)**
Ln(exchange rate)	0.829	0.829	0.821	0.822
	(18.78)**	(18.68)**	(18.68)**	(18.67)**
Ln(absorption)	0.676	0.677	0.671	0.669
	(18.89)**	(18.85)**	(18.89)**	(18.79)**
Ln(1+tariff)	-3.981	-4.019	-3.948	-3.919
	(13.16)**	(13.12)**	(12.91)**	(12.89)**
Ln(distance)	-0.782	-0.781	-0.775	-0.775
	(34.44)**	(34.50)**	(34.40)**	(34.24)**
Ln(unit costs)	-0.565	-0.659	-0.590	-0.525
	(14.40)**	(16.14)**	(15.10)**	(13.88)**
Ln(CR)*R&D	0.031			0.025
	(3.15)**			(2.48)*
R&D	0.198			0.170
	(2.84)**			(2.43)*
Ln(CR)*non-family		0.009		0.002
		(0.26)		(0.08)
Non-family dummy		0.407		0.246
		(1.54)		(1.02)
Ln(CR)*MNC			0.053	0.043
			(2.03)*	(1.62)
MNC dummy			1.150	0.983
			(4.88)**	(4.21)**
Fixed Effects	Industry-Year	Industry-Year	Industry-Year	Industry-Year
Estimation Method	Heckit	Heckit	Heckit	Heckit
Observations	64256	64256	64256	64256

Notes: Table displays coefficients and t-statistics for marginal effects obtained via Heckit (t-statistics based on standard errors clustered on industry-destination years). The dependent variable is firm-destination exports. ME denotes “marginal effect”. Marginal effects are evaluated at sample means. +, * and ** signify statistical significance at the 10%, 5% and 1% levels.

Table 6: Counterfactual Experiments I

Counterfactual	Annualized counterfactual change in aggregate export growth rate (%)			
	Heckman (Industry-Year FE)			Poisson (Industry-Year FE)
	"Intensive"	"Extensive"	Total	
Absorption unchanged	-3.31%	-0.93%	-3.99%	-5.06%
- EU15 only	-1.20%	-0.14%	-1.33%	-1.99%
- Europe_other only	-0.72%	-0.22%	-0.85%	-1.13%
- NAFTA only	-0.17%	-0.05%	-0.22%	-0.32%
Unit labor costs unchanged	1.18%	0.09%	1.37%	0.99%
Exchange rates unchanged	1.46%	0.26%	2.00%	2.63%
- C&S America only	0.69%	0.18%	1.01%	1.27%
- All except C&S America	0.70%	0.06%	0.90%	1.32%
Tariffs unchanged	0.21%	0.14%	0.34%	0.14%
CR unchanged	0.12%	0.04%	0.17%	0.34%
- No. exporters	0.04%	0.00%	0.04%	0.17%
- Exchange rates	-0.01%	-0.00%	-0.01%	-0.02%
- Tariffs	0.03%	0.01%	0.04%	0.07%
- Unit labor costs	0.06%	0.03%	0.08%	0.11%

Notes: Table reports annualized differences in growth rates between the counterfactual scenario indicated in the first column and actual export growth rates. Results are based on coefficient estimates obtained via the estimation method indicated at the top of each column. See text for details.

Table 7: Counterfactual Experiments II

Counterfactual	Annualized counterfactual change in aggregate export growth rate			
	Heckman (Industry-Year FE)			Poisson (Industry-Year FE)
	"Intensive"	"Extensive"	Total	
Chinese counterfactuals	-0.13%	-0.10%	-0.20%	-0.18%
- Absorption growth	-0.09%	-0.06%	-0.13%	-0.14%
- Market crowding	0.01%	0.00%	0.02%	0.03%
- Chinese import tariffs	-0.04%	-0.05%	-0.09%	-0.05%
- EUR/RMB exch. rate	-0.01%	-0.01%	-0.01%	-0.01%
Higher absorption growth EU15	0.45%	0.05%	0.51%	0.77%

Notes: Table reports annualized differences in growth rates between the counterfactual scenario indicated in the first column and actual export growth rates. Results are based on coefficient estimates obtained via the estimation method indicated at the top of each column. See text for details.