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The international elasticity puzzle is worse than you think☆

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ABSTRACT

We instrument export prices with firm level electricity cost shocks and estimate three international price elasticities using firm-level export data: the elasticity of firm exports to export price, tariff and real exchange rate shocks. In standard models these three elasticities should be equal. We find that this is far from being the case. The export price elasticity is the highest, around 5, much larger than the exchange rate elasticity. The international elasticity puzzle is therefore worse than previously thought. We also show that exporters absorb one third of tariff changes in their export prices. Because we take into account this reaction of export prices to tariffs, our estimate of the tariff elasticity corrects from this omitted-variable bias.

1. Introduction

The elasticity of substitution between Home and Foreign varieties, the Armington elasticity, is a crucial parameter in international trade and macroeconomic models. It is a fundamental primitive that shapes the international transmission of shocks into prices and quantities, and also a key component for analyzing the welfare impacts of trade liberalization (Arkolakis et al., 2012).[1] However, no consensus has emerged on its value and a tension between the micro and macro views on this elasticity exists. The evidence suggests that the elasticity of export volumes to changes in tariffs is quite large (typically above 2)whereas the elasticity to changes in exchange rates is small (typically around one or lower). This is what Ruhl (2008) has dubbed the international elasticity puzzle. As shown by previous studies, the elasticity puzzle is not only observed with macroeconomic or sectoral data but also with firm level data (Fitzgerald and Haller (2018)).

Our paper contributes to this literature by putting firm level export prices explicitly at the center of the analysis of the international elasticity puzzle. The elasticity puzzle literature has focused mainly on the difference in elasticities between tariffs and exchange rates but has not considered the elasticity of firm level export volumes - at the intensive margin - to firm level export prices. This is so even though in standard trade models these three elasticities should be identical.[2] Indeed, if a firm i of country h exports goods to country j, trade models with CES preferences predict the export volume to be:

$$Export_{i,h,j} = \left(\frac{p^{x}_{ih,j}\tau_{h,j}}{\varepsilon_{h,j}}\right)^{-\sigma} P_j^{\sigma-1} Y_j \quad (1)$$

In addition to the destination appropriate price index P_j and income Y_j the firm level export volumes depend on $p^{x}_{ih,j}$, $\varepsilon_{h,j}$ and $\tau_{h,j}$ the export price in domestic currency, the bilateral exchange rate and the tariff respectively. If the destination price index P_j (which could be affected by individual price, tariff and exchange rate movements) is controlled for, the (intensive margin) elasticity of exports to these three variables should be identical and equal to σ. We show this is far from being the

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[1] Arkolakis et al. (2012) show that for a larger class of trade models, the welfare gain from trade can be expressed as a function of the change in the share of domestic expenditure and the trade elasticity to variable trade costs.

[2] See Krugman (1979), Melitz (2003), Eaton and Kortum (2002) or Atkeson and Burstein (2008)

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0022-1996/

case in the data. Moreover, previous empirical studies, by omitting changes in firm level export prices, implicitly assume that these do not react to exchange rate and tariff shocks. We show this assumption is not valid as exporters absorb one third of tariff changes in their prices. This generates an omitted-variable bias in the estimates of the elasticity of firm level exports to tariffs and exchange rates.

An obvious difficulty to estimate the export price elasticity is that export prices and export quantities are endogenous at the firm level. This problem does not occur for exchange rates and tariffs shocks that could be considered exogenous to a single firm. To overcome this difficulty we use a firm level time varying instrumental variable for export prices. The instrumental variable we choose comes from an original dataset providing information on a firm specific cost shock, namely firm level electricity prices.

We argue below that these firm level electricity cost shocks are related to factors exogenous to its export performance (regulation changes, year and length of beginning of contracts, national and local tax changes, location, changes in both market and regulated prices and local weather) and are likely to affect its export performance only through the firm export price. We match this dataset to a dataset on French export volumes and values to estimate the firm level export price elasticity. We do this by using French exporters data on the period 1996–2010, we focus on the intensive margin of trade and find an export price elasticity around 5.

By introducing explicitly firm level export prices as a determinant of export volumes we also improve on the estimation of the elasticity of exports to exchange rate and tariff shocks. This is because we take into account the reaction of export prices to exchange rate and tariff shocks to estimate the elasticity of exports to those shocks and therefore eliminate an omitted-variable bias in existing studies. This would not be important if exporters did not absorb part of the tariff or exchange rate changes in their FOB domestic currency export price. However, we find that French exporters reduce substantially (by 3.5%) their export price when faced with a tariff increase of 10%. This is much less so for exchange rate movements. This implies that if export prices are not controlled for, the estimated elasticity of export volumes to tariff and exchange rate movements is the addition of two opposite channels: the true firm-level elasticity of exports to these shocks and the export elasticity to the endogenous reaction of export prices to exchange rates and tariffs movements.

Our results confirm that, when estimated at the firm level, the tariff elasticity is higher than the exchange rate elasticity. This is the standard international elasticity puzzle. We go further by showing that the difference between the export price and the exchange rate elasticities is even larger. In this sense, the international elasticity puzzle is worse than previously thought.

Our paper is related to a large literature that has estimated the elasticity of exports to tariffs and exchange rates. An extensive review of the literature estimating trade elasticities is provided by Hillberry and Hummels (2012). Fitzgerald and Haller (2018) and Berman et al. (2012) found that the elasticity of a firm export volumes to an exchange rate movement was below unity and around 0.5 to 0.7, close to our own estimate. The impact of those shocks on export volumes typically depends on how exporters pass them into export prices, how importers pass them into consumer prices and how final consumers react to change in final goods prices. In addition, Fitzgerald and Haller (2018) find that the elasticity of firm level exports to tariffs is larger than to exchange rate movements, the international elasticity puzzle at the firm level, which we confirm on French data. This elasticity also depends on the extent of strategic complementarities between firms in price setting, an issue analyzed by Amiti et al., 2016 and which we take into account in our analysis. Amiti et al. (2016) also estimate the price response to a firm specific cost shock (approximated with changes in the unit values of the imported intermediate inputs) but do not analyze the response of exports to these cost shocks. As in our paper, Piveteau and Smagghue (2015) use an instrumental strategy to estimate export price elasticities which they find to be above 2.[3] They use exchange rate variations interacted with firm-specific importing shares as instruments and concentrate their analysis on the estimate of quality of exported goods. We argue that an advantage of using electricity cost shocks as instruments rather than exchange rates is that they are more likely to affect exports only through their impact on export prices.

On the tariff side, using product-level information on trade flows and tariffs, Head and Ries (2001), Romalis (2007) and Caliendo and Parro (2015) estimate aggregate elasticities of 6.9, 8.5 and 4.5 respectively. Also using sector-level data, Costinot et al. (2012) find an elasticity of 6.5. Aggregate elasticities are also estimated by Imbs and Mejean (2015) and Berthou and Fontagné (2016). Finally, Anderson and van Wincoop (2003) survey the evidence on the elasticity of demand for imports at the sectoral level and conclude that this elasticity is likely to be in the range of 5 to 10. Using the methodology of Feenstra (1994), Broda and Weinstein (2006) report median import demand elasticities of 3.7 on their most disaggregated samples. Simonovska and Waugh (2014a) estimate the aggregate elasticity to be around 4 and Simonovska and Waugh (2014b) show how to estimate aggregate trade elasticities using price gaps. Our paper differs fundamentally from those using aggregate (sector-level) data in terms of both objective and interpretation. Given that the elasticity to tariffs we estimate is at the firm level, this means that its interpretation is different from the elasticity of aggregate trade to tariffs present in models such as Eaton and Kortum (2002), where the aggregate elasticity of trade to tariffs is governed by the heterogeneity across goods in countries' relative efficiencies and not the elasticity of substitution between domestic and foreign goods. Papers that provide estimates of trade elasticity with firm-level trade data are closer to our contribution. Buono and Lalanne (2012) exploit the Uruguay Round phasing out of tariffs and estimate a tariff elasticity for French firms at the sector level close to 2. Fitzgerald and Haller (2018) also exploit the time dimension of this Round and obtain a tariff elasticity around 3 for exports of Irish firms.[4] Berthou and Fontagné (2016) use firm-level export data for France, control for firm-level unobserved characteristics and multilateral resistance terms in each destination, and estimate a mean elasticity of the product-destinationfirm-level exports with respect to applied tariffs around 2.5.[5] These firm-level estimates are close to our own tariff elasticity estimate around 2.

Bas et al. (2017) use an alternative and innovative methodology to identify the firm-level trade elasticity with respect to applied tariffs; i.e. the differential treatment of exporters from two distinct countries (France and China) in a set of product-destination markets in a single year. One advantage of their identification strategy, compared to the existing literature, is that it does not rely on a single origin country and thus it fully controls for multilateral resistance terms. With an identification strategy based on cross-country variation in tariffs, they obtain an elasticity around 5, higher than ours and higher than the other firm-level estimates but close to our export price elasticity. In comparison, our identification strategy, like other authors already cited, relies on changes over time in bilateral tariffs. The dimension of the identification appears to be the main reason for the difference in estimates.

Our paper is also related to Feenstra, Luck, Obstfeld and Russ (2014) who distinguish between the elasticity governing the substitution between home and foreign goods (which they call macro and estimate to be below 1) and the elasticity governing the substitution between varieties of foreign goods (which they call micro and estimate around 4.4). Our approach is different as: (i) we do not make this distinction; (ii) we

[3] Erkel-Rousse and Mirza (2002) also use instruments for aggregate import prices (exchange rates, wage rates and production and exporter fixed effects) to estimate trade elasticities and find that they vary between −4 and −15.

[4] We refer to the coefficient −3.2 in their table 13: this is the export revenue (firm-product level) response to tariff shocks over the 1997–2009 period for Irish firms exporting to the 30 largest Irish exports destinations and for more than 6 years.

[5] The transformation of the exports variable relies on a benchmark destination – the United States– for the firm and a reference population of firms exporting to that destination.

use exporter level data rather than sectoral data on imports and (iii) we rely on an instrumentation method (firm level electricity cost shocks) rather than a GMM estimator that rests on the assumption that demand and supply costs are unrelated. This assumption may be an issue if higher costs of production are correlated with higher quality.[6] A further advantage of our instrument is that it bypasses the problem of quality that may affect both demand and supply costs. Indeed, an electricity price change in one year is plausibly uncorrelated with a quality change on the exported product in that year.

Finally our paper is related to the pass-through literature (see Burstein and Gopinath (2014) and Bussiere et al. (2016) for a survey). Our first stage results show that French export prices absorb much more tariff than exchange rate shocks. In particular, a 10% increase in tariff implies a 3.5% reduction in the export price; while a 10% depreciation of the bilateral real exchange rate causes a 0.2% increase in the export price.

The rest of the paper is structured as follows. We present data and our instrumental variable for export prices in Section 1. Our results on the estimate of the elasticity of export volumes to (instrumented) export prices are given in Section 2. We then estimate jointly and compare the elasticities of exports to export prices, tariffs and exchange rates in Section 3. The last section concludes and attempts to provide an interpretation of our results in particular relating trade elasticities to the volatility and persistence of shocks.

2. Data and instrumental variable description

2.1. Data

We use three confidential firm level datasets: (i)*Douanes* database on French firms exports, (ii) *Ficus/Fare* on French firms balance sheet information and (iii) *EACEI* data on energy consumption and purchase of French firms.[7] Macro level control variables come from standard sources (World Bank, CEPII and Penn World Table).

The *Douanes* database is provided by French customs for the period 1995–2010 on import and export flows of French firms by destination country, product (HS 6-digit classification) and year. This database contains all trade flows by firm-product-destination that are above 1000 euros for extra EU trade and 39,000 euros for intra-EU trade, so it can be considered an exhaustive sample of all French exporting firms.[8] Based on export values and quantity (reported in kilos) we compute the Trade Unit Values (TUV) for a specific firm-product (HS 6-digit)-destination-year cell (here used as proxy for the export price). So in what follows the export price $p_{i,j}^{x}$ is always approximated by TUV. The potential amount of observations is thus very large: there are almost 100,000 exporting firms per year and 200 destination markets. For this reason (and also because our main instrumental variable does not vary with product dimension - see below), we collapse the French customs data at firm-destination-year level. Hence, the resulting TUV is the weighted average across exported products of a given firm-destination-year cell.[9] By doing so, we lose the HS exported product dimension; but when needed, we still have the information of the main *industry* (NAF700 classification) in which the firm operates (as coded by the INSEE).[10]

The weighted average of TUVs can suffer from a composition bias (due to the aggregation of several products exported within a firm-destination-year cell).[11] In Section 2.2 we provide a detailed discussion on the potential aggregation bias and how we address it. We argue that using the core product of the firm helps to reduce this bias concern. In these robustness tests, for each firm-destination we keep the HS-6 code that represents the maximum (average across years) exported value for the firm-destination. Thus, in all the core product estimations we refer to a specific *sector* rather than to a more general industry dimension (as done in the baseline dataset described above).

The second firm level database is *Ficus/Fare* which contains balance sheet information for all French firms. From this database we use employment level of each French manufacturing firm as a control variable in our main regressions. From *Ficus/Fare* we also keep the labor cost and the purchase of intermediate inputs and raw materials used to compute the share of electricity over the total cost reported below.

The information on firm level electricity price (used as instrumental variable for the export price, see Section 2) is provided by the *EACEI* survey on energy purchase and consumption by around 11.000 French firms in the period 1996–2010.[12] For each plant-year combination we have information about the use of different types of energy such as electricity, steam, coke and gas. For consistency with the French custom data, the *EACEI* database has been aggregated at firm level by summing electricity bill and consumption across plants within the same firm.[13] The price of electricity has been computed as the ratio between electricity bill (in) and purchased quantity of electricity (in kWh). The final electricity price for the firm is thus expressed in €/kWh. When we merge the three firm level databases we are left with around 8500 exporters per year.

Finally we merge firm level data with other macro datasets: (i) OECD.stat for the GDP of destination countries, (ii) CEPII MacMap HS-4 and HS-6 data for tariffs and (iii) Penn World Table for nominal exchange rates and consumer price indexes (used to calculate the real exchange rate). The MacMap database on tariffs records ad-valorem*applied* tariff for each country pair-sector (HS-4 and HS-6 digit) observed in four years: 2001, 2004, 2007 and 2010 (see Bouet et al. (2008) and Guimbard et al., 2012for more details on MacMap).[14] Since French exporters do not face tariff in EU, we simply set to zero intra-EU tariffs. As described above, for our baseline regressions we use a firm-destination-year specific dataset. Accordingly we need firm-destination-year specific tariff, which has been computed as the weighted average tariff faced by a firm into a given destination-year (average across exported products).[15] In the core product estimations, since we keep the core product exported by each firm, we can use the (core) product level tariff. In the main regressions, we use HS4-digit tariffs. This has the advantage that it reduces the concern that tariffs may be targeted on some French firms which experience high export growth and therefore that tariffs may not be fully exogenous at the firm level. In a robustness check, we use tariffs at the HS6-digit product level.

2.2. Firm level electricity prices as instruments for export prices

In our empirical strategy, we use the firm specific electricity price as an instrumental variable for the export price.[16] The average electricity price in our dataset (reported in Table 1) is in line with the publicly available average prices for the manufacturing sector. Importantly, our dataset exhibits variance across firms and within a firm over time. We

6 See Feenstra and Romalis (2014) for how taking into account the issue of endogenous quality alters the estimation of international price elasticities.

7 All firm level confidential dataset have been used at CEPII.

8 Reporting of firms having trade values below such thresholds is discretionary but many firms below the bar are in the dataset.

9 We use export quantities as weights.

10 Notice that each firm is assigned to a unique industry of activity by INSEE. There are 615 industries in the NAF700 classification.

11 The majority of firm-destination cells (60%) involve export shipments within a unique HS4-digit heading. Since products within a HS4-digit headings are mostly homogeneous, the composition bias concern is reduced in our context.

12 The survey has been conducted on firms with more than 20 employees.

13 We use the French firm identifier *siren* to merge with the Custom database.

14 We use tariff in 2001 for the years preceding 2001. Tariffs in 2001 were also used for the period 2001–2003. Then tariffs in 2004 have been used for the period 2004–2007. Finally, tariffs in 2007 were used for tariffs in the period 2007–2010. As a robustness check we also use tariff data from WITS (World Bank).

15 We follow Berthou and Fontagné (2016) and use the product share over firm's total exports, all destinations, as a weight.

16 Ganapati, Shapiro and Walker (2016) use energy cost shocks as instruments for marginal cost shocks. Their aim, very different from ours, is to estimate the pass-through of those shocks into domestic prices. A major difference with our paper is that they use the interaction between national fuel prices for electricity generation and 5-year lagged electricity generation shares at the state level. We use firm level data for electricity prices.

Table 1
In-sample descriptive statistics on firm-year level dataset.

	Mean	Std dev	Min	Max	Std dev between	Std dev within
Electricity Price (€/kwh)	0.064	0.016	0.033	0.139	0.016	0.009
Electricity cost share	0.027	0.059	0.000	0.999	0.059	0.043
Trade Unit Value (ln)	2.256	1.673	−1.660	7.982	1.667	0.479

Source: Author's calculation on Ficus/Fare sample and Douane data.

also observe annual variations in prices that are not synchronized across firms. In Fig. 1 the dotted line is the average price of electricity paid by French firms between 1996 and 2010. We also show the price paid by two anonymous firms chosen here to have a mean price and a standard deviation similar to the mean and the standard deviation of the overall sample. Although the overall trend is similar (first downward then upward), we see that these firms experience yearly shocks that are very different.

We now move to explain what there is behind the firm specific component of electricity prices in the French manufacturing sector. In particular we argue how the specificities of the French electricity market enable us to use firm level electricity prices as an instrument for export prices. Note that our regressions will include firm fixed effects so that any time invariant characteristic of the firm electricity price will be controlled for and that the source of variation we will use is across years for a given firm. A characteristic of the French electricity market is that many contracts co-exist with both regulated and market driven prices. Regulated prices are offered only by EDF (the main historical operator) and unregulated prices are offered by all operators to all firms (Alterna, Direct Energie, EDF, Enercoop, GDF Suez, Poweo, and others). Firms can also have several contracts with several producers, and some firms may also produce their own electricity.

Another characteristic is that many firms had to renegotiate long-term contracts that ended during the period. These long term contracts allowed firms to have lower prices and their expiration means that firms may experience an increase in price in different years depending on the year the contract was initially signed and its length. Importantly for us there has also been many changes in regulations during the period 2001–2010. Under the pressure of the European Commission the market has been partially deregulated and opened with an increasing role of both imports and exports. Large firms were the first to be able to opt out from regulated prices in 2000 and this possibility was open progressively to all firms in the 2000s. However, in the same period many different electricity tariffs co-existed and were affected by several

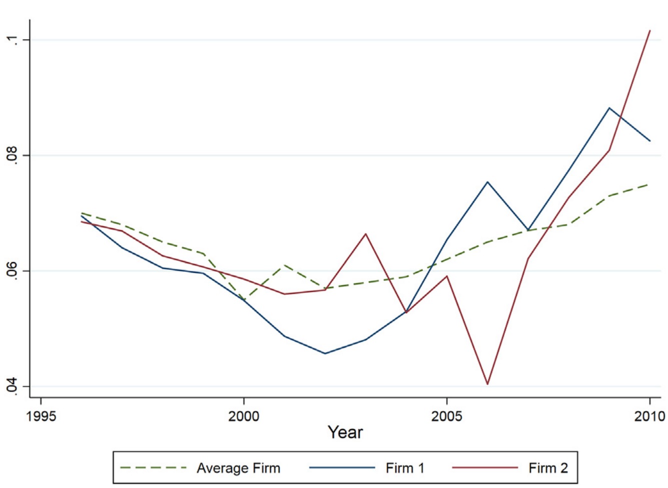

Fig. 1. Electricity Price (€/kwh) over the period 1996–2010. Average and two specific firms. Note: The dashed line refers to the average firm, obtained by collapsing the dataset by year. Firm 1 and 2 are specific (anonymous) firms having mean and std. dev electricity price similar to sample mean and std. dev. Source: Authors based on EACEI dataset.

changes. For example, in 2006 there was a large increase in electricity prices for firms that had opted (in the preceding years) for contracts with deregulated market prices. The government decided in 2007 to allow those firms to go back to a transitory regulated tariff (TarTAM tariff) calculated on the basis of the regulated tariff plus a surcharge depending on the firm of 10%, 20% or 23%. Not all firms chose to do so as it depended on the difference between the firm specific previous contracted price and the (firm specific) TarTAM (transitory regulated tariff). This choice depended itself on the date the previous contract was signed. This possibility was then stopped in particular because it was deemed to be a sectoral subsidy by the European Commission and this meant another change in price for some but not all firms. There are also different regulated tariffs for firms. The Blue tariff (small electricity users) allows a fixed price (for a year) with possibility to have lower prices during the night. Yellow and Green tariffs (intermediate and large electricity users) may also benefit from a fixed price with lower average prices during the year if they accept to pay higher prices possibly on a maximum 22 days in the year (very cold days in winter when household demand is high). Depending on the location of the firm in France these price increases may differ. Also, some firms benefit from low prices because they are close to hydroelectric facilities. Finally, the electricity price also depends on several taxes especially the so-calledTURPE (to pay for distribution and transport in particular) since 2000 which was created after the European Commission obliged France to separate the production and the distribution of electricity. The tax is itself quite complex, firm specific (in particular it is reduced if the firm has experienced a power outage of more than 6 h in the year) and changes every year. It can constitute up to 40% of the final electricity cost. Another tax (CSPE to finance renewable costs) also varies every year. Finally there are additional taxes at the city and department level that can vary both across locations and years.

This description of the electricity market in France shows that electricity prices vary at the firm level for reasons that are both endogenous to the firm activity (in particular its *average* electricity use, which is then captured by firm fixed effects in our empirical strategy) and more importantly exogenous to the firm export activity (regulation changes, year and length of beginning of contract, tax changes both at the national and local levels, location, changes in both market and regulated tariffs, local weather).[17] We take into account some of the impact of firm characteristics on electricity prices by including a firm fixed effect as well as a time varying measure of its activity (employment). Using firm specific electricity price changes as an instrument for export prices in the regression to estimate the price elasticity of exports is also valid because we believe that changes in the electricity price at the firm level affect export volumes only through their effect on export prices (the exclusion restriction). This would not be the case for other types of costs (wages or intermediate inputs) that may alter export volumes if an increase in these costs is caused by an increase in the quality of

[17] One may be concerned that the electricity price of French firms is correlated with the international price of energy which itself affects world demand for French exports. International price for gas and oil are indeed correlated with the French electricity price (the correlation is around 0.5 on the period). However, we include year fixed effects in all our regressions (either as year fixed effects alone or as destination-year fixed effects). This fully controls for the business cycle and world energy prices that might drive both the electricity prices in France and aggregate exports. Given the presence of firm and year fixed effects, this means our IV uses the part of firm level electricity prices which is not common to all French firms and which deviates from the average over the period for this firm.

the good (see Piveteau and Smagghue, 2015 on this). An alternative instrument for marginal cost shocks is exchange rate shocks for intermediate imported inputs as in Piveteau and Smagghue (2015) and Loecker and Biesebroeck (2016). One potential issue with using the exchange rate variations (even if interacted with firm-specific importing shares) as an instrument is that they may affect directly the exports of the firm other than through its export price.

Our first stage regression that we detail in the next section resembles a pass-through equation where export prices depend in particular on electricity prices. In a standard framework where a firm i uses several inputs (electricity among others) which are imperfect substitutes and minimizes costs, the path-through of a firm level electricity cost shock p_{ei} to export prices p_i is given by:

$$\frac{dp_i}{dp_{ei}}\frac{p_{ei}}{p_i} = \frac{p_{ei}e_i}{p_{ei}e_i + \sum_{m=1}^{M} p_m x_{mi}} \tag{2}$$

where M is the number of inputs (other than electricity) and $p_m x_{mi}$ the expenditures on those inputs. Hence, the pass-through of electricity cost shocks to export prices is simply the share of electricity costs in the total costs of the firm. For each firm we have labor costs, energy costs and intermediate goods costs but not capital costs. In our data set which is restricted to the manufacturing sector this ratio is around 2.7% (see Table 1) so we should expect that in our first stage regressions the pass-through of a firm level electricity price shock to export prices is around the same number.[18] An alternative instrument for the firm specific export price, consistent with Eq. (2), would be the interaction between the firm-year specific price of electricity and the firm specific share of electricity over total costs. For this cost share we tried either the average share for the firm on the whole period or the share for the sector to reduce endogeneity. The advantage of this instrument is that it uses an information specific to the firm or the sector which describes its electricity intensity. The disadvantage is that total costs (including labor costs and intermediates) may be endogenous to exports of the firm, and it may affect the export of the firm through the mix of produced (and then exported) goods. We use this alternative instrument in robustness checks in Section 3.2 and find similar results.

The share of electricity over the total cost (as reported in Table 1) is computed as the ratio between the electricity bill and the total production costs of the firms available in the Ficus/Fare dataset (i.e. labor cost, purchase of intermediate inputs, raw materials and electricity). Table 2 reports the summary statistics for the sample of firms we use in our baseline regressions, so the number of firms and the other statistics reported in the table refer to a sample of exporting firms for which we also have balance sheet and electricity bill data. The average size of the firm over the period 1996–2010 is large but this is not surprising since these are exporting firms only.[19] There is also some variation in the electricity cost share over time: from 1.9% in 2005 up to 3.6% in 2002 and back to 2.5% in 2010 (the average over the period is 2.7%).

Our empirical strategy proceeds in two steps. First, we estimate the elasticity of export volumes to prices by using an instrumental variable approach to solve the endogeneity problem of prices. Then, we analyze the international elasticity puzzle in our data set by including in the same regression export price (instrumented), real exchange rate and firm specific tariffs.

3. Export volumes elasticity to export prices

To estimate the elasticity of export volumes to export prices we use the instrumental variable described in the previous section. We first report the baseline results with several combinations of fixed effects and controls and then report several robustness checks: using the core product of the firm, controlling for strategic complementarity and using various subsets of the data.

Table 2
In-sample summary statistics.

Year	N. firms	Employees	Elec. price	Elec. share
1996	9000	227	0.070	0.029
1997	9492	217	0.068	0.029
1998	9746	215	0.065	0.028
1999	9702	213	0.063	0.028
2000	5561	289	0.055	0.020
2001	8744	223	0.061	0.025
2002	5895	344	0.057	0.036
2003	5715	353	0.058	0.036
2004	6054	316	0.059	0.035
2005	4613	241	0.062	0.019
2006	6198	205	0.065	0.020
2007	6464	201	0.067	0.022
2008	5413	223	0.068	0.021
2009	5437	194	0.073	0.033
2010	5721	183	0.075	0.025

Notes: statistics on the sample of firms used in the baseline estimations.
Source: Authors' calculations on EACEI and Douane dataset.

3.1. Baseline results

The second stage regression has the following econometric specification depending on the set of fixed effects included:

$$ln\left(Export_{i,j,t}\right) = \theta_i + \theta_{jt} + \sigma ln\left(p^x_{i,j,t}\right) + \beta_1 ln\left(Emplo_{i,t}\right) + \varepsilon_{i,j,t} \tag{3}$$

$$ln\left(Export_{i,j,t}\right) = \theta_i + \theta_{jst} + \sigma ln\left(p^x_{i,j,t}\right) + \beta_1 ln\left(Emplo_{i,t}\right) + \varepsilon_{i,j,t} \tag{4}$$

$$ln\left(Export_{i,j,t}\right) = \theta_{ij} + \theta_t + \sigma ln\left(p^x_{i,j,t}\right) + \beta_1 ln\left(Emplo_{i,t}\right) + \beta_2\left(\mathbf{Z_j}.\mathbf{t}\right) + \varepsilon_{i,j,t} \tag{5}$$

while the first stage regression is the following respectively with the electricity price ($Elect.\ Price_{i,\ t}$) as instrument variable:

$$ln\left(p^x_{i,j,t}\right) = \theta_i + \theta_{jt} + \gamma_1 ln\left(Elect.Price_{i,t}\right) + \gamma_2 ln\left(Emplo_{i,t}\right) + \eta_{i,j,t} \tag{6}$$

$$ln\left(p^x_{i,j,t}\right) = \theta_i + \theta_{jst} + \gamma_1 ln\left(Elect.Price_{i,t}\right) + \gamma_2 ln\left(Emplo_{i,t}\right) + \eta_{i,j,t} \tag{7}$$

$$ln\left(p^x_{i,j,t}\right) = \theta_{ij} + \theta_t + \gamma_1 ln\left(Elect.Price_{i,t}\right) + \gamma_2 ln\left(Emplo_{i,t}\right) + \gamma_3\left(\mathbf{Z_j}.\mathbf{t}\right) + \eta_{i,j,t} \tag{8}$$

where subscripts i,j, s and t stand respectively for firm, destination market, sector and year. The dependent variable in Eqs. (3), (4) and (5) ln ($Export_{i,\ j,\ t}$) is the log of the exported volume by firm i in a specific country j and year t.[20] As stated in the introduction, a trade model with CES preferences predicts export volumes at the firmlevel of the form: $Export_{i,j,t} = \left(\frac{p^x_{i,j,t} \times tariff_{i,j,t}}{RER_{j,t}}\right)^{-\sigma} P^{\sigma-1}_{j,t} Y_{j,t}$.[21] In log terms, this expression produces a testable equation of the form givenin Eqs. (3), (4) and (5) with the export price ($p^x_{i,\ j,\ t}$) as our main variable of interest. Tariffs and real

[18] The observed distribution of the electricity cost share among French manufacturing exporters is shown in appendix figure A1 and suggests that although heterogeneous this cost share is concentrated around its mean.

[19] Moreover, remember that EACEI survey is conducted on firms with more than 20 employees.

[20] Note that contrary to the majority of trade elasticity estimates, we use export volumes rather than values. The reason is that we explicitly introduce instrumented prices. We also ran our regressions using the value of exports and our results are indeed robust to this transformation of the left-hand side variable, the difference in the elasticity being indeed 1.

[21] With respect to equation (1) we add the subscript t indicating the time dimension, and remove the subscript h indicating the country of firm i because in the empirics we use only French firms.

exchange rates are added in the next section. As for the destination country aggregate income and price index, they are either captured by destination-year fixed effects (in regressions 3 and 4) or by GDP and real effective exchange rate of country j in regression (5)- they enter in the vector of controls $\mathbf{Z_{j,t}}$. Independently of the structure of fixed effects, we include the employment of the firm i at time t (in log) as a control for the size of the firm -$ln(Emplo_{i,\ t})$. Our main focus here is the instrumented log of the export price (i.e. trade unit value) -$ln(p^{x}_{i,\ j,\ t})$ - and we expect a negative coefficient for σ. As explained in the data section we use two main regression samples: (i) exported volumes and average export price across products within firm-destination-year (*baseline full dataset*), (ii) exported volumes and export price of the HS-6 specific core product of the firm for a given destination (*core product dataset*). The subscript s refers to industry (NAF700 classification) and sector (HS classification) respectively for baseline and core product dataset (see section 1 for further details).

We compare our estimates of the price elasticity with various fixed effect combinations. In Eqs. (3) and (6), we include firm fixed effects (θ_i) and destination-year fixed effects (θ_{jt}). This enables to control for any time invariant characteristic of the firm and for any destination specific time varying impact on the firm demand, as well as components of the global cycle that may affect electricity prices and French exports. Firm fixed effects control for the time invariant part of the firm characteristics (e.g. quality) that may affect its exported volumes. Destination-year fixed effects control for the effect of the macroeconomic cycle in the destination country (aggregate demand) as well as the destination price index (see Anderson and van Wincoop (2003) and Head and Mayer (2014)). This set of fixed effects is standard in the trade literature. In eq. (4) and (7) the specification is more demanding as it replaces destination-year fixed effects (θ_{jt}) by sector-destination-year fixed effects (θ_{jst}). In the core product estimations, when s represents the sector of firm's export, the θ_{jst} fixed effects should better control for both the aggregate demand for the sector and the price index at destination as it takes into account differences across sectors in a same destination-year cell. Moreover, θ_{jst} fixed effects control for sector specific shock in each destination.[22] In Eqs. (5) and (8) we include firm-destination (θ_{ij}) and year (θ_t) fixed effects. These fixed effects properly control for any time shock (common to all destinations) and for any firm-destination specific characteristics affecting the export volumes of French firms. Because the specification in Eq. (5) does not control for the destination time varying aggregate demand and price index through the proper fixed effects, we add a set of country-year specific variables $\mathbf{Z_{j,t}}$ including GDP (in ln) and effective real exchange rate as a proxy for the destination price index (computed as in Berman et al. (2012)) (Tables A1-A3).

Table 3 shows the results of the IV regression with the three different sets of fixed effects and the first stage results at the bottom of the table. The coefficient on electricity prices is always positive and significant, and the F-stat is always above 15. Note in particular that the first stage estimates of the impact of electricity cost shocks on export prices are very stable as they vary between 0.04 and 0.05.[23] As discussed before, a simple model predicts that this elasticity should be close to the share of electricity costs in total costs. The average observed share in our sample is around 3% so not very different from our estimated pass-through.

Table 3 provides a first estimate of the export price elasticity that varies (in absolute value) between 3.9 and 5.7.[24] In the specification reported in columns 1 of Table 3, firm fixed effects and destination year fixed effects are included but there is no control for the time varying activity of the firm that may affect electricity prices, export prices and export volumes. This is added in column 2 where we control for the employment of the firm (in ln). Alternatively, we also used firm turnover as a firm-year control and obtained similar results. However, we prefer to include employment as turnover includes also export sales (i. e. dependent variable). Note, that our first stage estimation is not much affected by this control. Then, in columns 3 and 4, the destination-year fixed effect is replaced by a more demanding sector-destination-year fixed effect.[25] Finally specifications 5 and 6 have a firm-destination fixed effect and a year fixed effect. Results are very similar.[26] All in all, we conclude that the estimate of the export price elasticity is robust across different specifications (i.e. fixed effects) and around 5.

3.2. Robustness checks addressing composition biases

The empirical strategy described in the previous section aims at estimating the export price elasticity for a single firm exporting to a given destination - see Eq.(1). For this reason we aggregate French custom data at firm-destination-year level by summing export volumes across products within a specific firm-destination-year cell. This aggregation - natural given that our IV is firm-year specific and does not vary across products - may imply two potential biases in estimation. First, if firms export different products towards a given destination, our estimation would produce an (unweighted) export price elasticity across the different products exported by the firm. Second, by aggregating across products, changes in unit values and export quantities may reflect changes in the product mix instead of true price changes. For example, an electricity price increase may push firms to concentrate on high quality exported goods and therefore to change its mix of exports towards potentially low elasticity products.

We reduce these two potential aggregation biases as in Berman et al. (2012) by restricting our estimate to the *core* product observations. Namely, for each firm-destination we keep the HS-6 product that represents the maximum level of exports over the period 1996–2010 (averaged across years). By doing so, changes in export prices and volumes do not reflect changes in the product mix of the firm, and for each firm-destination we estimate a single elasticity rather than an average of elasticities across exported products. Results reported in column 1 in Table 4 show that, without sector-destination-year fixed effects, the estimated elasticity is higher (5.5 versus 3.9 in absolute value) than that obtained on the full sample (column 2 of Table 3 where we do not restrict to the core product of the firm). We acknowledge however that using the core product is an imperfect solution to the biases discussed above. The reason is that the core product being the best performing product, its elasticity may be lower than for other products. However, when sector-destination-year fixed effects are included, the aggregated and the core-product elasticities are almost identical (compare column 4 of Table 3 to column 2 in Table 4).[27] Therefore there is no strong suggestion that firms faced with a cost shock tilt their product mix towards higher quality, lower elasticity products. The F-stat of the first stage regressions are above the rule of thumb of 10 supporting the absence of weak instrument concern (this is further confirmed by Crag-DonaldF-stat above the critical value).

A second potential bias, coming from entry-and-exit dynamics of firms to a given destination, may affect our estimates. A positive shock to electricity costs and export prices may lead to exit of firms with weak unobserved idiosyncratic performance. Negative shocks would conversely lead to exit of weak firms. This composition or selection bias may therefore lead to underestimate the (absolute) value of the export price elasticity. We follow a similar method to Fitzgerald and Haller (2018) to reduce this bias by restricting the sample to firms with long export histories. The reason is that these firms should be further away

22 Sector-destination-year fixed effects in the baseline sample estimations use the NAF700 classification of each firm, i.e. the sector to which the firm belongs to.

23 The full first stage regression results are shown in the appendix in Table A7.

24 We report the OLS estimation in the appendix in Table A2. Not surprisingly the demand elasticity is lower in absolute value, just above 1, when we do not instrument the export price. An obvious reason is that in this case price movements are not exogenous and affected by demand shocks to the firm.

25 Industries are defined using the NAF700 4-digit classification of the French statistical institute INSEE for each firm. There are 615 NAF700 industries in the French economy.

26 Another possible specification is to run the regression in first difference. The coefficient on $p^{x}_{i,\ j,\ t}$ in the second stage is estimated at −5.176 and significant. However, the instrument in the first stage although significant is weak with an F-stat less than 4.

27 Because we use the core product of the firm we can use the HS classification for the sector fixed effect. We use the 4 digit level of HS.

Table 3
Baseline 2SLS regressions on full dataset.

Dep var	Export volumes (ln)					
	(1)	(2)	(3)	(4)	(5)	(6)
$p^x_{i,j,t}$ (ln)	−4.203***	−3.916***	−5.692***	−5.366***	−5.544***	−5.131***
	(0.729)	(0.671)	(1.197)	(1.125)	(0.982)	(0.900)
Employment (ln)		0.159***		0.132***		0.205***
		(0.012)		(0.017)		(0.015)
GDP (ln)					0.784***	0.831***
					(0.167)	(0.153)
Effective RER (ln)					−0.067***	−0.067***
					(0.017)	(0.016)
Firm FE	Yes	Yes	Yes	Yes	No	No
Destination-Year FE	Yes	Yes	No	No	No	No
Firm-Destination FE	No	No	No	No	Yes	Yes
Year FE	No	No	No	No	Yes	Yes
Sector-Destination-Year FE	No	No	Yes	Yes	No	No
First Stage Coefficients						
Electricity Price	0.049***	0.050***	0.040***	0.040***	0.046***	0.046***
Employment (ln)		0.002		−0.002		−0.001
First Stage Statistics						
F-stat	23.25	23.47	15.83	15.60	22.83	22.67
R-squared	0.770	0.770	0.779	0.779	0.883	0.883
Cragg-Donald F-stat	66.91	67.51	42.44	41.82	82.44	81.53
Observations	1,630,856	1,630,856	1,630,856	1,630,856	1,488,954	1,488,954

Standard errors are clustered within firm-year in all estimations. When Destination-Sector-Year FE are included, the sector is the main NAF700 sector of the firm. More details on the first stage results are reported in Table A7. Critical value for the Cragg-Donald F-stat is 16.38.
*** $p < 0.01$;
** $p < 0.05$;
* $p < 0.1$

from the entry/exit thresholds at which a cost shock would lead to a change in participation choice. One possibility is to restrict the sample to core-product observations of firms exporting to a given destination over the entire period (1996–2010). As shown in columns 5–6 in Table 4, the sample is reduced drastically. The sample may also select firms that are very high productive and not representative. Hence, we prefer a less demanding sample, and in columns 3–4 in Table 4 we keep core product observations of firms exporting more than 5 years to a given destination. We think this is a good compromise (similar to the choice of Fitzgerald and Haller (2018)) to reduce the composition bias without relying on a small sample of very specific firms.

In this restricted sample (columns 3–4 in Table 4), the estimated elasticity is very similar to our main estimates in table (3)i.e. around 5. It is higher without sector-destination-year fixed effects (6.0 compared to 3.9 in the full dataset and to 5.5 in the full sample but core product). However, with sector-destination-year fixed effects, this estimate is very similar both to our baseline estimate and to our estimate using core product observations in the full sample (5.5 against 5.4). The

Table 4
Core product robustness checks.

Dep var	Export volumes (ln)					
	(1)	(2)	(3)	(4)	(5)	(6)
$p^x_{i,j,t}$ (ln)	−5.533***	−5.441***	−5.989***	−5.525***	−5.296***	−3.374***
	(1.261)	(1.126)	(1.425)	(1.228)	(1.667)	(0.803)
Employment (ln)	0.144***	0.127***	0.196***	0.148***	0.201***	0.171***
	(0.017)	(0.017)	(0.024)	(0.029)	(0.027)	(0.020)
Sample	Full		Exporting more than 5 years		Balanced	
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Destination-Year FE	Yes	No	Yes	No	Yes	No
Sector-Destination-Year FE	No	Yes	No	Yes	No	Yes
First Stage Coefficients						
Electricity Price	0.037***	0.039***	0.042***	0.044***	0.043***	0.065***
Employment (ln)	−0.001	−0.005	−0.004	−0.015***	−0.002	−0.006
First Stage Statistics						
F-stat	15.24	12.40	14.27	10.10	8.75	8.19
R-squared	0.831	0.910	0.848	0.924	0.893	0.956
Cragg-Donald F-stat	30.54	25.14	28.31	22.04	13.25	17.59
Observations	1,045,502	1,045,502	643,564	643,564	173,827	173,827

When Sector-Destination-Year FE are included, the sector is the HS 4-digit chapter of the core HS 6-digit product.
Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-Donald F-stat is 16.38.
*** $p < 0.01$;
** $p < 0.05$;
* $p < 0.1$

joint F-stat is lower (but still above 10) when using core products estimates for firms exporting more than 5 years. In columns 5–6 in Table 4, the robustness check using core product observations for firm exporting over the entire period shows that the export price elasticity remains qualitatively similar but with a moderate weak instrument issue which can be explained by a drastic fall in the number of observations. As explained above, this is the reason why our preferred robustness check is using the sample of firms exporting for more than 5 years. We acknowledge that this is an imperfect way to address the composition bias but we are reassured that as we reduce the bias, our estimates remain broadly unchanged. In what follows we adopt this sample of observations as a main robustness check.[28]

3.3. Robustness checks addressing strategic complementarity

A final issue we want to address is strategic complementarity that has been emphasized recently by Amiti et al. (2016) in international pricing. The concern is that in the first stage regression, the electricity cost shock that generates the export price increase could also lead close competitors to increase their own price. In turn, this may alter the estimate of the impact of the export price increase on its export sales. If such a strategic complementarity exists, for example of the kind analyzed by Atkeson and Burstein (2008), the perceived elasticity of demand is different (smaller) from the elasticity of substitution across products. A complete analysis of this issue is beyond the scope of our paper but we can take advantage of our dataset to check whether our estimates are robust to a crude measure of these strategic complementarities. Note that they should be already taken into account when we include sector-destination-year fixed effects as in columns 3 and 4 of Table 3, and/or columns 2, 4 and 6 in Table 4. The reason is that in a model such as Atkeson and Burstein (2008), strategic complementarities exist as long as firms are large enough to affect the sectoral price index. Asector-destination-year fixed effect should control for the sector price index and therefore for strategic complementarity. Our results are robust to the inclusion of sector-destination-year fixed effects with sectors defined either with NAF700 classification (Table 3) or HS4 classification (Table 4).

As a further robustness test, we use the core product dataset and control for the prices of other French exporters to the same destination and in the same HS6 sector. We follow the empirical strategy of Amiti et al. (2016) although we depart from them because we use a different instrumental variable and we analyze the strategic complementarity on export prices while they analyze it on domestic prices. We proceed in two steps. First we control for strategic complementarity of French exporters only and then we control for strategic complementarity of non-French exporters to the destination.

In order to define the relevant set of competitors, we need the specific HS6-digit in which the firm operates. So, for this estimations we rely on the core product based sample of firms (as in Table 4). For firm i exporting to a given HS6-destination combination, we define the French competitors export price (TUV) as the average price of French firms exporting to that HS6-destination combination. We exclude from this average the export price of firm i. We also exclude from the sample HS6-destination combinations with less than two competitors. Finally, we define foreign competitors TUV as the average import price (TUV) of non-French exporters to a given HS6-destination where the French firm i is exporting (using BACI dataset). In Table 5 we show results based on the core product dataset (core product of firms exporting more than 5 years) controlling for the average price of French competitors - column 1 - and for both domestic and foreign competitors' TUVs - column 3. The results are intuitive as firm export prices increase with both domestic and foreign competitors' prices (in the first stage) suggesting the presence of strategic complementarity. The price

Table 5
Strategic complementarity robustness checks.

Dep var	Export volumes (ln)		
	(1)	(2)	(3)
$p^x_{i,j,t}$ (ln)	−6.069***	−6.177***	−5.663***
	(1.631)	(1.662)	(1.488)
Employment (ln)	0.164***	0.159***	0.165***
	(0.041)	(0.043)	(0.039)
Export price competitors (ln)	0.597***		0.498***
	(0.206)		(0.175)
TUV importing country (ln)			0.205***
			(0.057)
Electricity Price competitors (ln)		0.236**	
		(0.093)	
Sample	Core product, exporting more than 5 years		
Firm FE	Yes	Yes	Yes
Destination-Year FE	Yes	Yes	Yes
First Stage Coefficients			
Electricity Price	0.046***	0.046***	0.047***
Employment (ln)	−0.017***	−0.017***	−0.018***
Export price competitors (ln)	0.127***		0.117***
TUV importing country (ln)			0.038***
Electricity Price competitors (ln)		0.005	
First Stage Statistics			
F-stat	11.96	11.86	12.60
R-squared	0.855	0.852	0.855
Cragg-Donald F-stat	18.30	18.07	19.27
Observations	301,746	301,746	298,333

Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-Donald F-stat is 16.38.
*** $p < 0.01$;
** $p < 0.05$;
* $p < 0.1$.

of competitors also have a positive impact on export volumes. However, the main result is that the estimated elasticity is not much affected.

As a robustness check in columns 2 of Table 5 we replace the French competitors prices by an exogenous shock to these prices, i.e. the average electricity cost for these French competitors. Its effect on export volumes is positive and significant in column 2 but again the estimate of the export price elasticity is not much affected. All in all, from this first set of evidence we conclude that our estimate of the firm level export price elasticity is precisely estimated and relatively high at around 5.

3.4. Other robustness checks

In the appendix Table A4 we report the estimate of the export price elasticity when dropping the top-5% and bottom-5% TUV firms within each HS6. Dropping these extreme values helps alleviate the concern that export prices are highly volatile and subject to measurement errors. Our results are robust to this new sample. As discussed in Section 1, the electricity market in France was deregulated in 2000. As a robustness check in Table A5 we provide estimation results using the post-2000 period only. Our results remain similar. The value of the export price elasticity does not change with the set of destinations countries. In appendix Table A6 we show a robustness check splitting the sample into EU vs non-EU destinations and the export price elasticity remains qualitatively unchanged.

4. Export elasticity to prices, tariffs and real exchange rates

In this section we compare the elasticity to the firm specific export price with two other trade elasticities often estimated in the existing literature: the elasticity (i)to tariff and (iii) to real exchange rate. Several issues arise in this comparison. First, we instrument export prices but not tariffs and exchange rates. This is maybe more a problem for the former than the later which can be thought as exogenous to the firm. Second, while cost shocks to export prices are firm-specific this is not the

[28] In Table A3 we report the share of total exports accounted by the core product of firms and by the core product of firms exported more than 5 years in a given destination. This is a substantial part (between around 60% and 50%) of total French exports.

case of exchange rate and tariff shocks. Finally, another limitation of an estimation of a tariff elasticity using firm level export data for a single origin country is that it cannot fully control for destination resistance terms. Thus the difference between the price and tariff elasticity should not be over-interpreted. The bottom line is that the export price elasticity (like the tariff elasticity) is much larger than the exchange rate elasticity, worsening the international elasticity puzzle.

4.1. Baseline results

In measuring how trade volumes react to tariffs and to real exchange rate movements, our main contribution is to take into account the fact that exporters can partly absorb tariff and exchange rate shocks in their export prices. If this is the case, existing firm-level estimates suffer from an omitted-variable bias. Our estimation strategy is the same as in Eq.(5) but we add firm-destination-year specific tariffs ($ln(tariff_{ijt}+1)$) and bilateral real exchange rate (RER_{jt}) as follows:

$$\begin{aligned} ln\left(Export_{i,j,t}\right) &= \theta_{ij} + \theta_t + \sigma_1 ln\left(p^x_{i,j,t}\right) + \sigma_2 ln(RER_{j,t}) \\ &\quad + \sigma_3 ln\left(tariff_{i,j,t}+1\right) \\ &\quad + \alpha_4 ln(Emplo_{i,t}) + \alpha_5(\mathbf{Z_{j,}t}) + \varepsilon_{i,j,t} \end{aligned} \tag{9}$$

while the first stage regression is:

$$\begin{aligned} ln\left(p^x_{i,j,t}\right) &= \theta_{ij} + \theta_t + \gamma_1 ln(Elect.Price_{i,t}) + \gamma_2 ln(RER_{j,t}) \\ &\quad + \gamma_3 ln\left(tariff_{i,j,t}+1\right) \\ &\quad + \gamma_4 ln(Emplo_{i,t}) + \gamma_5(\mathbf{Z_{j,}t}) + \eta_{i,j,t} \end{aligned} \tag{10}$$

All variables have the same meaning as before. In contrast to the specifications we tested in the previous section, we can only include firm-destination(θ_{ij}) and year (θ_t) fixed effects since destination-year fixed effects would be perfectly collinear with real exchange rates.[29] As before we include a set of destination-year specific control variables $\mathbf{Z_{jt}}$ containing the GDP (in log) of destination countries to control for import demand and the real *effective* exchange rate to control for the degree of competition in the destination country and the price index of the importing country. One advantage of including the (instrumented) price in the export volume equation is that it enables us to take into account that exporters absorb part of a change in tariff and exchange rate in their FOB export price in exporter's currency. The international elasticity of tariffs and exchange rates are actually a mix of two elasticities: the direct impact of tariffs on trade and the indirect impact through the change in export prices that absorb part of the tariff change. Hence, using the notation in Eqs.(9) and (10), the standard OLS elasticity that does not control for the export price would be $\sigma_3+\gamma_3\sigma_1<\sigma_3$ if $\gamma_3<0$.

The results are shown in Table 6. Column (1)is our preferred regression. The first stage regression can be interpreted as a pass-through equation and it is interesting to comment. We find that tariffs and real exchange rates shocks are partly absorbed by exporters in their export prices. Only a small part of the exchange rate change is absorbed (less than 3%). The pricing to market behavior is more relevant for core product sample estimations reported in table (7), where around 10% of the exchange rate shock is absorbed in the export price. This result is consistent with the evidence in Berman et al. (2012). Bussiere et al. (2016) find (on disaggregated bilateral trade flows) that the pass-through of exchange rate into export prices is in general larger for industrialized countries than for emerging ones. Our results on French exporters is consistent with this. The most interesting result is for tariffs. Indeed, to our knowledge, we are the first to report that exporters absorb a large part of tariff changes in their export prices: they indeed increase export prices by 3.5% following a 10% decrease in tariff.

Table 6
2SLS regressions on full baseline dataset.

	Dep var: export volumes (ln)			
	(1)	(2)	(3)	(4)
$p^x_{i,j,t}$	−5.171***	−5.434***	−4.681***	−5.504***
	(0.911)	(1.001)	(1.200)	(1.104)
RER (ln)	0.659***	0.831***	0.566***	0.636***
	(0.040)	(0.086)	(0.037)	(0.038)
Ln(tariff+1)	−1.771***	−2.508***	−1.534***	−1.891***
	(0.175)	(0.628)	(0.405)	(0.406)
Effective RER (ln)	0.121***	0.026	0.117***	0.089***
	(0.019)	(0.051)	(0.019)	(0.019)
Employment (ln)	0.205***	0.248***	0.144***	0.173***
	(0.015)	(0.017)	(0.021)	(0.018)
GDP (ln)	0.624***	0.530**	0.817***	0.702***
	(0.175)	(0.260)	(0.168)	(0.169)
Estimator	2SLS	2SLS	2SLS	2SLS
Sample	All countries	OECD	non-OECD	extra-EU
Firm-Destination FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
First Stage Coefficients				
Electricity Price	0.046***	0.049***	0.042***	0.045***
RER (ln)	0.026***	0.061***	0.006	0.009
Ln(tariff+1)	−0.348***	−0.498***	−0.327***	−0.353***
Effective RER (ln)	0.009	−0.009	0.007	0.001
Employment (ln)	−0.002	0.003	−0.009	−0.006
GDP (ln)	−0.181***	−0.235***	−0.131***	−0.142***
First Stage Statistics				
F-stat	22.47	23.08	9.10	16.53
R-squared	0.883	0.882	0.883	0.880
Cragg-Donald F-stat	80.87	56.44	25.96	48.45
Observations	1,488,954	863,035	625,919	1,022,174

Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-Donald F-stat is 16.38. More details on the first stage results for specification in column 1 is reported in table 15.
*** $p<0.01$;
** $p<0.05$;
* $p<0.1$.

We now comment the second stage results. Table 6 shows that the inclusion of tariffs and real exchange rates does not alter the estimate of the instrumented export price elasticity that remains close to 5. In regression (1), we report our main result on the ranking of the three elasticities: the export price elasticity is much larger than the (possibly underestimated) elasticity for the tariff which itself is much larger than the elasticity for the real exchange rate which is around 0.6.

In regressions reported in columns (2), (3)and (4)we analyze the difference between exports towards OECD, non OECD and extra-EU countries. We first note that the absorption of both exchange rates and tariffs in export prices is more relevant towards OECD countries. We also note that the coefficients in the second stage are relatively similar across destinations although slightly lower towards non OECD countries. This may suggest that French goods are less substitutable with non-OECD produced goods.

Controlling for the export price is crucial as tariffs have a direct impact and an indirect impact through export prices. An increase in tariffs affects export volumes via the elasticity of substitution effect (measured here as −1.77, see regression (1) in Table 6) and via the impact it has on export price which themselves affect export volumes ($-0.35^{*}-5.17=1.81$). The two effects in our data approximately cancel each other in our data and an OLS regression not controlling for export prices would lead to an estimated elasticity not significantly different from zero, although it would be clearly wrong to conclude from such a regression that tariffs do not affect firm level exports.[30]

[29] In a robustness check reported in Table 7 we exclude the real exchange rate from the sample of covariates and run a specification including destination-year fixed effects.

[30] Results are available upon request.

Table 7
Robustness checks.

	Dep var: export volumes (ln)								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$p^x_{i,j,t}$	−5.065***	−5.260***	−5.306**	−5.195***	−3.605***	−6.755***	−5.501***	−5.166***	−5.128***
	(0.864)	(1.070)	(2.390)	(0.1064)	(0.903)	(1.712)	(1.804)	(0.909)	(0.905)
RER (ln)		0.791***	0.552***	1.010***	1.148***	1.102***	1.060***	0.658***	0.653***
		(0.054)	(0.120)	(0.067)	(0.078)	(0.149)	(0.205)	(0.041)	(0.039)
Ln(tariff+1)	−2.116***	−2.000***	−2.070*	−1.509***	−1.629	−0.380**	−0.717***	−1.605***	−0.395***
	(0.475)	(0.457)	(1.100)	(0.294)	(0.351)	(0.162)	(0.180)	(0.308)	(0.082)
Effective RER (ln)		0.151***	−0.027	0.150***	0.173***	0.099***	0.130**	0.121***	0.130***
		(0.024)	(0.033)	(0.035)	(0.038)	(0.031)	(0.051)	(0.020)	(0.020)
Employment (ln)	0.207***	0.176***	0.111***	0.297***	0.352***	0.106***	0.194***	0.205***	0.206***
	(0.014)	(0.016)	(0.022)	(0.036)	(0.041)	(0.026)	(0.028)	(0.015)	(0.015)
GDP (ln)		0.897***	1.239***	1.543***	1.824***	0.701***	1.119***	0.634***	0.723***
		(0.165)	(0.306)	(0.134)	(0.145)	(0.240)	(0.174)	(0.173)	(0.157)
Estimator Sample	2SLS Full	2SLS Full 1996–2007	2SLS Full First Diff.	2SLS Full Top-25	2SLS Full Top-10	2SLS Core more 5 years	2SLS Core Balanced	2SLS Full Tariff HS6	2SLS Full Tariff WITS
Firm-Destination FE	Yes	Yes	No	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Destination-Year FE	Yes	No	No	No	No	No	No	No	No
First Stage Coefficients									
Electricity Price	0.052***	0.044***	0.015*	0.074***	0.094***	0.037***	0.041***	0.046***	0.047***
RER (ln)		0.035***	0.045***	0.033**	0.051**	0.082***	0.105***	0.026***	0.024***
Ln(tariff+1)	−0.228***	−0.409***	−0.453***	−0.168***	−0.244***	0.019	0.008	−0.318***	−0.040**
Effective RER (ln)		0.013**	0.004	0.013	0.017	0.002	−0.000	0.009*	0.011**
Employment (ln)	0.001	−0.001	0.006*	−0.017**	−0.026**	−0.006	−0.002	−0.002	−0.001
GDP (ln)		−0.139***	−0.116***	−0.094***	−0.111***	−0.131***	−0.069***	−0.179***	−0.161***
First Stage Statistics									
F-stat	24.04	16.87	3.6	15.20	10.49	12.77	8.13	22.52	22.44
R-squared	0.78	0.897	0.001	0.913	0.92	0.935	0.946	0.883	0.884
Cragg-Donald F-stat	84.94	57.86	6.96	77.32	53.29	44.29	23.21	81.04	81.66
Observations	1,496,270	1,218,470	1,003,361	403,196	167,273	640,447	172,918	1,488,954	1,459,931

Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-Donald F-stat is 16.38.
*** p < 0.01.
** p < 0.05.
* p < 0.1.

Another interesting result of Table 6 is the impact of cyclical demand in the destination country. French exporters *decrease* their export prices towards destinations where GDP is higher than average. The impact is substantial: a 1% increase in GDP at destination, leads French exporters to decrease their price to that destination by 0.18% (column 1 Table 6). This implies that the elasticity of exports to destination demand shock is the sum of two components: the direct effect of the foreign demand on exports (0.62 coefficient in the second stage regression in Table 6 column 1), and the effect of the fall in export prices when GDP at destination increase ($-0.18^{*}-5.2=0.94$).[31] This suggests that more than half of the increase in exports following an increase in the GDP at destination, is due to the firms' pricing strategy rather than to a standard foreign demand effect. To our knowledge, our paper is the first to document this pricing behavior which cannot be directly reconciled with existing models. Although Atkeson and Burstein (2008) may be a good candidate, we leave to future research the aim to rationalize this pricing strategy.[32]

4.2. Robustness checks

In Table 7 we perform a series of robustness tests. In column (1), we replace year fixed effects by destination-year fixed effects. In this case, the exchange rate variable is absorbed by the fixed effect but the impact of the tariff which is firm-destination specific can still be estimated. The destination-year fixed effect enables to better control for the impact of changes in tariffs on the destination price index. The estimated elasticity for the export price and the tariff remain similar to our preferred specification in Table 6. In column (2), we restrict the sample to pre-crisis years (1996–2007) and the results are robust. In column 3, we run the regression in first difference (with control variables and IV both expressed in first difference). The estimated coefficients on the three elasticities are similar but the instrument is weak.

An empirical concern is the selection bias in the export status if firms select endogenously in different destinations. In heterogeneous firm trade models, only high-productivity firms are able to serve more distant and more costly markets. In our framework, low productivity firms will exit destinations with higher tariff or depreciated exchange rates and this selection effect may bias our elasticity. To address this problem, we follow Fitzgerald and Haller (2018) and Mulligan and Rubinstein (2008) and run robustness checks using a subsample of top exporting firms, i.e. firms with exports above 75^{th} and 90^{th} percentile of destination-sector specific distribution of exports. Results, reported in columns (4)and (5)of Table 7 suggest that the selection bias is not a big problem in our data. Our results are also robust on the estimate of the export price elasticity to using the core-product for both firms exporting more than five years (column 6)and for firms exporting over the entire period (columns 7). As discussed above, this partially solves the aggregation and selection biases.

[31] This is the combination between the first stage GDP coefficient on export price and the second stage export price on export volumes.

[32] Atkeson and Burstein (2008) show in a model of imperfect competition and variable markups that because firms market shares determine the price elasticity of demand, firms decrease their markups and prices when they lose market share. If a destination GDP boom attracts new firms and products and therefore reduces the French exporters' market share, this would increase the elasticity of demand and provide an incentive to reduce export prices. A related but different mechanism is introduced by Jaravel (2016) who shows that increasing market size causes the introduction of more products and, through increased competitive pressure and decreasing markups, lower prices. This is also coherent with a Melitz and Ottaviano (2008) type of model where, departing from CES assumptions, competition is tougher (markups lower) in larger markets accommodating more firms.

In columns (8)and (9), we use alternative data sets for tariffs, respectively HS-6 classification and WITS data set (HS4). These robustness checks answer the concern that measurement issues may explain the observed difference in the tariff and electricity-instrumented own price elasticity. Imprecision in the measurement of tariffs may indeed lead to an attenuation bias driving the tariff elasticity towards zero. The tariff elasticity (−1.605) in column (8)is similar to the one obtained in Table 6 column (1). Another potential measurement issue is the data source used for tariffs; substituting WITS to MAcMaps in column (9)of Table 7 reduces the tariff elasticity estimate but not the export price elasticity.

Finally, in Table A8, we test the robustness of our results to the use of an alternative instrument for firm export prices. Consistent with Eq. (2), we use the interaction between the firm-year specific electricity price and the share of electricity costs in the total firm's costs. For this cost share we use either the average share for the firm on the whole period or the share for the sector to reduce endogeneity. The advantage of this instrument is that it uses a specific information about the firm or sector specific electricity intensity (as in Eq. (2)). The drawback of this approach is the likely endogeneity bias. Indeed, the total cost of the firms (in particular labor costs and intermediates) may be endogenous to the export performance of the firm and may affect exports through other channels than the export price. In particular, the mix of exported goods might be affected by labor and/or intermediates cost if such products are labor or intermediate goods intensive. Hence, we think this instrument may be more reliable when used in the core product sample. The results are shown in Table A8. The instrument works well in the sense that in the first stage the elasticity of export price to the instrument is between 0.8 and 0.9 (regressions 1 and 2) in the full sample. This elasticity should be around unity if there was full pass-through of costs to prices. The international elasticities are similar to those estimated with our main instrument except that they are smaller especially for tariffs and export prices. As explained above however, we believe that this instrument is more reliable for the core product sample which is shown in regressions (3)and (4). In this case the first stage is weak. The estimated elasticities are very similar to those estimated with our main instrument (see Table 6) for the export price and the exchange rate but lower for the tariff.

5. Conclusion and interpretation

The main contribution of this paper is to offer an estimate of the firm level price elasticity of exports using an original instrumental variable strategy. Our results point robustly to an estimate around 5. The second contribution points to the importance of the absorption of tariff changes by exporters in their export prices. This implies that an estimate of elasticities of exports to tariffs that does not take into account the endogenous reaction of export prices is a mix of two opposite effects: the elasticity of substitution between home and foreign goods and the elasticity of exports to the endogenous reaction of export prices to the tariff shock. These two effects have opposite signs: an increase in tariff generates a substitution away from French exports but the endogenous fall in French exporter prices counteracts this. Our third contribution is to show that the price (or tariff) elasticity is much higher in absolute value than the exchange rate (around 0.6) elasticity: the international elasticity puzzle is well alive and actually worse than previously thought.

Our results can be viewed as stylized facts in search of theory. We now briefly present two interpretations that are consistent with our

Table 8
Ranking of elasticities, volatilities and persistence.

	Exchange rate	Tariff	Export price
Trade elasticity	0.6	2	5
Coeff. var. (firm-destination)	7.8%	0.9%	0.5%
Persistence	67%	88%	95%

Note: The coefficient of variation reported in this table is the simple average across firm-destination specific coefficients of variation measures. The persistence measure reported in the last row is the AR(1)coefficient from destination-year specific regressions for each of the variables reported in the first row. Country fixed effects are included in the AR(1)regressions: $y_{j,\,t} = \beta y_{j,\,t-1} + \theta_j + \varepsilon_j t$.

data. One interpretation of our results, which we cannot test, is that importers and wholesalers in the destination country absorb differently in their prices these three shocks or that they switch to alternative producers differently depending on the source of the shock. If these intermediaries pass export price and tariff shocks to retail prices more than exchange rate shocks this could explain the ranking we observe. This could in turn be due to differences in the perceived persistence and volatility of those shocks. Consistent with the model of Drozd and Nosal (2012), importers and retailers absorb more volatile and less persistent shocks because they need to explicitly build market shares by matching with their customers. If this process is costly and time consuming, it may be that they will do it only when shocks are not too volatile and persistent enough. We cannot properly test this mechanism because we do not observe import and consumer prices of the exported French goods. However, we can compare the ranking of our elasticities to the ranking of volatilities and persistence for the three shocks. We calculated in our sample the coefficients of variation for the three shocks. To be consistent with the dimension of our estimation, we calculated it for each firm-destination and then computed the average. For export prices, we estimate the volatility that comes from our specific cost shock. Hence, we take the predicted value from the first stage of the estimation reported in column (4)of Table 6 but exclude tariffs and exchange rate. For persistence, we estimated the coefficient on an AR(1) process.

As shown in Table 8, a larger volatility and lower persistence are associated with a lower elasticity. Although only suggestive this comparison points to an interpretation of differences of the international elasticity to the three shocks as being linked to their differences in their respective volatility and persistence.

Another avenue to interpret our results, in particular on the difference in the elasticity of trade with respect to tariffs or exchange rates on the one hand and to export prices on the other hand is that the first two affect all Eurozone exporters (for tariffs all EU exporters) whereas a change in export price is firm specific. In the first case, following a tariff reduction or a euro depreciation, the demand for transport, distribution and marketing services in the destination will increase (see Arkolakis (2010)) which may push up costs in the destination. Importers, wholesalers and retailers experiencing an increase in costs following an expansion on French exports, may pass-through these costs to consumers so that the fall in consumer prices may be smal. This would not (or at a much lower extent) be the case with a single French firm export expansion following a decrease of its export price. The pass-through to consumer prices would therefore be smaller in the case of a shock that affects all French exporters than in the case of a firm specific cost shock. In turn, this implies a lower trade elasticity for a tariff, or exchange rate shock than for a firm specific shock. Our results are consistent with such a mechanism. We leave to further research to explore this possible explanation of the international elasticity puzzle (Fig. A1).

Appendix A. Appendix

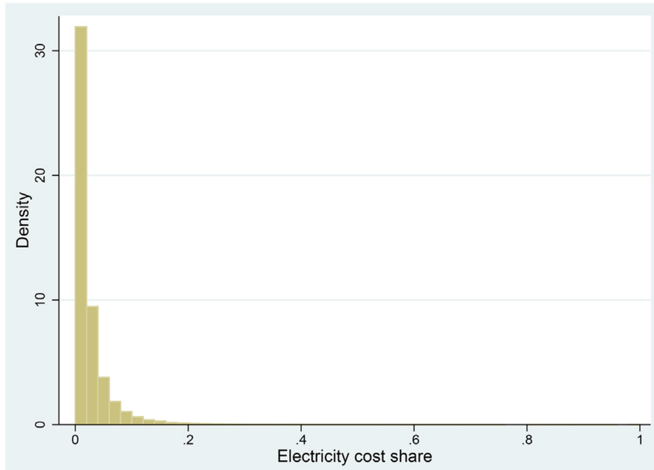

Fig. A1. Histogram of electricity cost shares.

Table A1
In-sample descriptive statistics before the firm-year aggregation.

	Observations	Mean	Std dev	Min	Max
Electricity Price (€/kwh)	1,630,856	0.062	0.015	0.033	0.139
Exported Quantity (ln)	1,630,856	8.378	3.187	−0.693	20.702
p_{ij}^{x} (ln)	1,630,856	2.608	1.813	−1.66	8.005
Employment (ln)	1,630,856	5.372	1.068	0.693	8.869
Turnover (ln)	1,630,856	10.407	1.471	−1.881	17.23
Ln (tariff+1)	1,630,856	0.042	0.084	0	2.397
RER (ln)	1,630,856	0.106	0.191	−2.005	1.162
GDP (ln)	1,630,856	26.05	1.925	18.3	30.24
Effective RER (ln)	1,630,856	1.179	1.967	−2.09	9.499

Table A2
OLS regressions on full baseline dataset.

Dep var	Export volumes (ln)	
	(1)	(2)
$p_{i,j,t}^{x}$ (ln)	−1.268***	−1.143***
	(0.003)	(0.002)
Employment (ln)	0.153***	0.215***
	(0.007)	(0.008)
GDP (ln)		1.457***
		(0.027)
Effective RER (ln)		−0.080***
		(0.008)
Firm FE	Yes	No
Destination-Year FE	Yes	No
Firm-Destination FE	No	Yes
Year FE	No	Yes
Observations	1,366,037	1,624,300
R-squared	0.621	0.873

Standard errors are clustered within firm-year in all estimations.
*** p<0.01;
** p < 0.05;
* p < 0.1.

Table A3
Sample size characteristics of core product firms: full sample vs firms exporting more than 5 years.

Sample	Median size	N. of firms	Share of total exports
Core product full sample	200	18,774	61%
Core product exporting more than 5 years	230	6763	49%

Table A4
Robustness check dropping top-5 and bottom-5TUV firms within HS6.

Dep var	Export volumes (ln)	
	(1)	(2)
$p^x_{i,j,t}$ (ln)	−4.143***	−5.733***
	(0.737)	(1.145)
Employment (ln)	0.154***	0.202***
	(0.021)	(0.016)
GDP (ln)		0.895***
		(0.163)
Effective RER (ln)		−0.041**
		(0.018)
Sample	Drop Top-5 and Bottom-5TUV firms within HS6	
Firm FE	Yes	No
Destination-Year FE	Yes	No
Firm-Destination FE	No	Yes
Year FE	No	Yes
First Stage Coefficients		
Electricity Price	0.047***	0.040***
Employment (ln)	0.000	−0.002
GDP (ln)		−0.131***
Effective RER (ln)		0.007*
First Stage Statistics		
F-stat	21.95	17.75
R-squared	0.789	0.894
Cragg-DonaldF-stat	65.20	67.22
Observations	1,554,061	1,424,017

Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-DonaldF-stat is 16.38.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

Table A5
Robustness check using post-2000 period.

Dep var	Export volumes (ln)	
	(1)	(2)
$p^x_{i,j,t}$ (ln)	−2.842***	−4.328***
	(0.586)	(1.086)
Employment (ln)	0.122***	0.185***
	(0.015)	(0.020)
GDP (ln)		0.792***
		(0.233)
Effective RER (ln)		−0.079***
		(0.233)
Sample	Period 2001–2007	
Firm FE	Yes	No
Destination-Year FE	Yes	No
Destination-Year FE	No	Yes
Year FE	No	Yes
First Stage Coefficients		
Electricity Price	0.052***	0.039***
Employment (ln)	0.017***	0.011**
GDP (ln)		−0.203***
Effective RER (ln)		−0.002
First Stage Statistics		
F-stat	18.09	11.48
R-squared	0.764	0.896
Cragg-DonaldF-stat	42.93	36.01
Observations	996,115	881,787

Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-DonaldF-stat is 16.38.
*** $p < 0.01$.
** $p < 0.05$;
* $p < 0.1$.

Table A6
Robustness check using intra-EU and extra-EU destination.

Dep Var	Export volumes (ln)		Export volumes (ln)	
	(1)	(2)	(3)	(4)
$p^{x}_{i,j,t}$ (ln)	−4.995***	−5.994***	−4.423***	−4.323***
	(1.036)	(1.220)	(0.910)	(0.909)
Employment (ln)	0.125***	0.158***	0.233***	0.271***
	(0.018)	(0.021)	(0.022)	(0.018)
GDP (ln)		0.974***		0.955***
		(0.146)		(0.120)
Effective RER (ln)		−0.080***		0.031
		(0.019)		(0.090)
Sample	Intra-EU		Extra-EU	
Firm FE	Yes	No	Yes	No
Destination-Year FE	Yes	No	Yes	No
Firm-Destination FE	No	Yes	No	Yes
Year FE	No	Yes	No	Yes
First Stage Coefficients				
Electricity Price	0.047***	0.046***	0.055***	0.051***
Employment (ln)	−0.003	−0.006	0.015***	0.011**
GDP (ln)		−0.106***		−0.062**
Effective RER (ln)		−0.001*		0.053**
First Stage Statistics				
F-stat	15.09	16.15	18.93	17.97
R-squared	0.764	0.883	0.811	0.888
Cragg-DonaldF-stat 35.05	45.42	40.33	43.09	
Observations	1,034,956	933,433	594,003	549,106

Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-DonaldF-stat is 16.38.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

Table A7
First stage regression results on full baseline dataset.

Dep var	$p^{x}_{i,j,t}$ (ln)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Electricity Price (ln)	0.049***	0.050***	0.040***	0.040***	0.046***	0.046***	0.046***
	(0.011)	(0.010)	(0.010)	(0.010)	(0.010)	(0.011)	(0.010)
RER (ln)							0.026***
							(0.008)
Ln(tariff+1)							−0.348***
							(0.029)
Employment (ln)		0.002		−0.002		−0.001	−0.002
		(0.004)		(0.004)		(0.004)	(0.004)
GDP (ln)					−0.158***	−0.158***	−0.181***
					(0.013)	(0.013)	(0.004)
Effective RER (ln)					0.003	0.003	0.009*
					(0.004)	(0.004)	(0.004)
Firm FE	Yes	Yes	Yes	Yes	No	No	No
Destination-Year FE	Yes	Yes	No	No	No	No	No
Firm-Destination FE	No	No	No	No	Yes	Yes	Yes
Year FE	No	No	No	No	Yes	Yes	Yes
Sector-Destination-Year FE	No	No	Yes	Yes	No	No	No
R-squared	0.770	0.770	0.779	0.779	0.883	0.883	0.883
F-stat	23.25	23.47	15.83	15.60	22.83	22.67	22.47
Cragg-DonaldF-stat	66.91	67.51	42.44	41.82	82.44	81.53	80.87
Observations	1,630,856	1,630,856	1,630,856	1,630,856	1,485,547	1,488,954	1,488,954

Standard errors are clustered within firm-year in all estimations. When Sector-Destination-YearFE are included, the sector is the main NAF700 sector of the firm. Critical value for the Cragg-DonaldF-stat is 16.38.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

Table A8
Robustness check using alternative instrumental variables.

Dep var	$p^x_{i,j,t}$ (ln)			
	(1)	(2)	(3)	(4)
$p^x_{i,j,t}$ (ln)	−2.378***	−1.297***	−4.117***	−3.620***
	(0.473)	(0.441)	(1.401)	(0.745)
RER (ln)	0.586***	0.557***	0.884***	0.843***
	(0.021)	(0.018)	(0.119)	(0.067)
Ln(tariff+1)	−0.799***	−0.423***	−0.430***	−0.439***
	(0.172)	(0.159)	(0.106)	(0.094)
Effective RER (ln)	0.095***	0.086***	0.093***	0.092***
	(0.011)	(0.008)	(0.019)	(0.017)
Employment (ln)	0.211***	0.214***	0.223***	0.226***
	(0.008)	(0.007)	(0.017)	(0.014)
GDP (ln)	1.127***	1.322***	1.044***	1.109***
	(0.089)	(0.082)	(0.186)	(0.106)
Sample	Full Sample		Core, exporting more than 5 years	
Firm-Destination FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
First Stage Coefficients				
Electricity Price*Avg. Elec Dependency	0.868***		0.768***	
Electricity Price*Sector Elec Dependency		0.806***		1.375***
RER (in log)	0.026***	0.026***	0.082***	0.082***
Ln(tariff+1)	−0.348***	−0.348***	0.019	0.021
Effective RER (ln)	0.009*	0.009*	0.002	0.002
Employment (ln)	0.002	0.002	−0.006	−0.006
GDP (ln)	−0.179***	−0.179***	−0.128***	−0.129***
First Stage Statistics				
F-stat	18.99	21.20	6.83	18.93
R-squared	0.883	0.883	0.935	0.935
Cragg-DonaldF-stat	51.41	84.62	29.85	68.97
Observations	1,488,954	1,488,954	640,447	640,447

Standard errors are clustered within firm-year in all estimations. Critical value for the Cragg-DonaldF-stat is 16.38.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.jinteco.2018.08.011.

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