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# Markup Heterogeneity, Export Status and the Establishment of the Euro

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**Abstract**

We investigate the effects of the establishment of the euro on the markups of French manufacturing firms. Merging firm-level census data with customs data, we estimate time-varying firm-specific markups and distinguish between eurozone exporters from other firms between 1995 and 2007. We find that the establishment of the euro has had a pronounced pro-competitive impact by reducing firm markups by 14 percentage points. By reducing export costs, the euro represented an opportunity for eurozone exporters to increase their margins relative to other firms. Quantile regressions show that the euro has led to a reduction in the variance of markups.

`Key Words`: Markups; Heterogeneity; Euro; Competition; Export Destinations.

`JEL codes`: C5; D43; F61; L16; L60.

## 1 Introduction

Surprisingly, the question of the pro-competitive effect of the establishment of the euro on markups has not been specifically addressed. The issue of the euro and its effect on trade and productivity has largely been discussed, beginning in 2000 with the so-called Rose effect (Rose 2000), which described the effect of a common currency on trade. As of today, however, the issue of the launch of the euro and its effect on firm market power has not been investigated *at the firm level*. This paper intends to do so in the case of French manufacturing firms.

Fifteen years after the launch of the euro, the financial crisis has challenged all positive arguments in favor of the euro and highlighted the inconsistencies of the eurozone: the lack of innovative monetary policy; the

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absence of an optimal currency area; and a lack of economic convergence. By considering price-cost margins, this paper returns to the main positive argument advocated by the proponents of the monetary union: the expected pro-competitive effect on consumer prices, a process thought to improve welfare considerably.

We expect the establishment of a unique currency to increase competition and therefore to decrease markups for three reasons. First, when all prices are expressed in the same currency, arbitrage by consumers is facilitated considerably. Firms that were previously sheltered by their domestic currency must then face competition with an array of foreign firms. The second reason is that the establishment of a single currency zone is similar to an increase in market size, thereby augmenting the number of firms in the market. Finally, the creation of the euro increased the profit opportunities associated with the access to a larger market, thereby intensifying firm entry into the market and increasing competition.

The key to our analysis is the distinction between 1999 and 2002, the two crucial dates concerning the new currency. In 1999, the launch of the euro as the new currency of account led to a decrease in export costs for eurozone exporters. Our intuition is that the establishment of fixed exchange rates in 1999 created an opportunity for eurozone exporters to increase their margins. Hence, far from being a homogeneous shock that was common to all firms, the launch of the euro concealed substantial heterogeneity in the response of firms. Depending on the extent to which firms passed through the cost cut to price or, alternatively, increased markups, the 1999 euro shock may have been an opportunity for a rise in markups. In 2002, the generalization of the use of the unique currency should have increased competition for the reasons mentioned above. Moreover, with stiffer competition in the eurozone, larger market size and new entries by firm are likely to have depressed firms' markups irrespective of the export status of the firm. This sequence in the introduction of the euro helps us to identify heterogenous responses by firms to the euro shock.

Our research is motivated by the use of a recent methodology to estimate markups, price over marginal cost, which varies by year and by firm. Until recently, markup estimates were constant across firms within an industry or a time period, generally one year (Hall 1986, Roeger 1995, Klette 1999). In contrast, the methodology developed by De Loecker & Warzynski (2012) generates time-varying, firm-specific markups. The key advantage is that this methodology allows us to investigate the asymmetrical effects of the euro on markups for different categories of firms, primarily types of exporters. Another advantage in the use of this method is that it accounts for changes in higher moments of the distribution of market power between firms, notably, their variance documenting heterogeneity in firm markups. Our understanding is that knowledge of the distribution is of crucial importance for the efficiency of any policy that aims to restore price-cost margins and their subsequent investments in quality upgrading and firm growth.

Whereas markups are expected to diminish after the establishment of the euro, their variance is also expected to decrease for two reasons. First, theoretically, firm markups have a lower bound at unity. This implies that the magnitude of adjustment for firms with initially low markups is narrow, whereas firms with higher markups can trim their margins more extensively. Hence, we hypothesise that a common currency shock should be more detrimental to firm markups with initially large market power. Second, starting with the seminal paper of Krugman (1987), a large literature has evidenced pricing-to-market behavior consisting of adjusting markups to changes in the exchange rate of the destination market. A partial pass-through of exchange rate has been evidenced in many countries, mainly in sectors that produce differentiated products. However, a partial pass-through has an immediate counterpart: a variation in markup per destination market. Therefore, the introduction of the euro should have reduced the variance in markup, reflecting the drop in the number of foreign currencies at stake.

Our empirical protocol uses the French census to gather financial information on manufacturing firms covering the 1994-2007 period with data on the export behavior of firms in terms of destination countries. The main specification is similar to implementing a difference-in-difference model and distinguishes firms that export exclusively to the eurozone from other exporters. This model accounts for serial correlation in the error term and for firm entry into and firm exit from domestic and foreign markets. We also develop two additional specifications. First, the dichotomous nature of dummy variables used in a difference-in-difference model makes no use of the amount of export to the eurozone. However, such continuous information should reflect the exposure of the firm to the euro shock. A model that accounts for this exposure is developed to allow firm markup to adapt gradually with the share of exports to the eurozone. Second, we rely on quantile regressions to account for the differentiated effect of the launch of the euro on firms, depending on their initial level of markups.

In a nutshell, we obtain three main results. First, we find that the generalization of the use of the euro in 2002 had a more pronounced procompetitive impact on markup – almost three times as large – as the establishment in 1999 of fixed exchange rates among eurozone countries. Second, the launch of the euro has been an asymmetric shock for French manufacturing firms. By reducing transaction costs for firms that trade with euro countries, the euro represented an opportunity for eurozone exporters to increase their price-cost margins relative to other firms in the sample. Third, the creation of the euro has led to a reduction in the heterogeneity of firm markups. As such, the establishment of the euro acted as a strong heterogeneity-reducing shock on markups.

This paper contributes to the controversial debate on the pros and cons of the euro. Past empirical literature has investigated the effect of the establishment of the euro on a series of economic phenomena. Initially documented by Rose (2000), the trade effect of the common currency was largely and intensely debated during the last decade, identifying a positive yet limited impact on trade[1]. Bun & Klaassen (2007) evidenced a lower bound amounting to 3%[2].

Regarding the introduction of the euro and its impact on firm markups, there is, to our knowledge, no empirical evidence using firm-level data. Cook (2011) provides an empirical evaluation of the impact of the euro at the macro level of markups. He shows a robust decrease in output labor share per industry in 21 OECD countries when a fixed exchange rate policy is implemented. He also finds a negative impact of eurozone membership on the share of income paid to workers. He interprets this increase as an increase in markups. However, this is an interpretation based on the fixed price hypothesis; nothing is said about productivity, which also affects the markups. At the firm level, Martin & Mejean (2012) document systematic changes in pricing strategies – not markups *per se* – related to the common currency and find that "monetary integration has heterogeneous effects on firms of different sizes". This paper is thus the first to focus on firm-specific markups and the impact of the introduction of the euro.

The paper is structured as follows. Section 2 provides detailed information on the method by which firm-level markups and productivity are simultaneously estimated. It also presents the database used and discusses preliminary evidence on the dynamics of markups, productivity, and export behavior in French manufacturing before and after the establishment of the common currency. Section 3 reports the results of the three econometric specifications: difference-in-difference, exposure to the euro shock, and quantile regression. Section 4 concludes.

# 2 Markup Estimates and Export Status

## 2.1 Simultaneous Estimations of Markups and Productivity

Similar to Hall (1986) and Roeger (1995), De Loecker & Warzynski (2012) rely on the production function framework. Unlike previous contributions, however, this framework neither imposes constant returns to scale nor requires the computation of the user cost of capital, a task that is difficult to perform accurately. Finally, this framework provides time-varying and firm-specific estimates of markups and productivity that allow us to unravel the heterogeneity in firms' markup.

[1]See Flam & Nordström (2006). For a survey, see Baldwin (2006).

[2]In France, Berthou & Fontagné (2008) found a small effect of the euro on trade stemming from the increase in the number of products exported per firm.

Let $Q$ be firm output as follows: $Q_{it} = Q_{it}(\mathbf{X}_{it}, K_{it})$, where subscripts $i$ and $t$ stand for firm $i$ at time $t$, respectively, $K$ is capital, and $\mathbf{X}$ is a vector of production factors. In this framework, capital is assumed to be fixed, whereas all remaining production factors are variable. We suppose that $Q(\cdot)$ is twice differentiable and continuous and that the objective of the producer is to minimize costs. The associated Lagrangian function then reads

$$\mathcal{L}_{it} = \mathbf{P}_{it}^{X}\mathbf{X}_{it} + r_{it}K_{it} + \lambda_{it}(\bar{Q_{it}} - Q_{it}(\mathbf{X}_{it}, K_{it})), \tag{1}$$

where $P_{it}^{X}$ and $r_{it}$ are firm input prices for input vector $\mathbf{X}$ and capital, respectively.

The first-order conditions satisfy

$$\frac{\partial \mathcal{L}_{it}}{\partial \mathbf{X}_{it}} = P_{it}^{X} - \lambda_{it}\frac{\partial Q_{it}(\mathbf{X}_{it}, K_{it})}{\partial \mathbf{X}_{it}} = 0 \tag{2}$$

and

$$\frac{\partial \mathcal{L}_{it}}{\partial Q_{it}} = \lambda_{it}, \tag{3}$$

which implies that $\lambda_{it}$ represents the marginal cost of production.

Rearranging (2) and multiplying both sides by $\frac{X_{it}}{Q_{it}}$ yields

$$\frac{\partial Q_{it}(\mathbf{X}_{it}, K_{it})}{\partial \mathbf{X}_{it}}\frac{\mathbf{X}_{it}}{Q_{it}} = \frac{P_{it}^{X}\mathbf{X}_{it}}{\lambda_{it}Q_{it}}. \tag{4}$$

The term on the left-hand side of Equation 4 is the output elasticity of the variable inputs $\mathbf{X}_{it}$, whereas the right-hand-side term is its share in total cost.[3] Now, defining firm markups $\mu$ as the price to marginal cost $\mu_{it} \equiv \frac{P_{it}}{\lambda it}$, it follows that $\lambda_{it} \equiv \frac{P_{it}}{\mu_{it}}$. Inserting the former into Equation 4 and simplifying yields

$$\mu_{it}^{X} = \frac{\theta_{it}^{X}}{\alpha_{it}^{X}}, \tag{5}$$

where the numerator $\theta_{it}^{X} = \frac{\partial Q_{it}(X_{it}, K_{it})}{\partial X_{it}}\frac{X_{it}}{Q_{it}}$ represents the output elasticity of input $\mathbf{X}_{it}$ and the denominator $\alpha_{it}^{X} = \frac{P_{it}^{X}\mathbf{X}_{it}}{P_{it}Q_{it}}$ is the share of input $\mathbf{X}_{it}$ in total sales. Hence, to compute the markup $\mu_{it}$, we need to compute both $\theta_{it}^{X}$ and $\alpha_{it}^{X}$ per firm and per time period. Although it is straightforward to compute $\alpha_{it}^{X}$, the estimation of $\theta_{it}^{X}$ is more demanding.

At the outset, two important choices need to be made explicit. First, we limit the set of variable inputs to labor $L$ only. One could think of inserting additional production factors, such as material $M$ or energy consumption

[3]This is true when at the optimal point of production, the marginal cost equalizes the average cost due to the free entry of firms into the market.

$E$. Theoretically, if all factor markets were perfect, the markup derived from material must yield the same value as the markup derived from labor: $\mu_{it}^M = \mu_{it}^L$. However, differences in factor markets' imperfections will yield different values of firm markups ($\mu_{it}^M \neq \mu_{it}^L$). Instead, we limit vector $\mathbf{X}$ to labor $L$ only and assume a perfect labor market[4]. This also implies that we define output $Q$ as value added.

The second important choice involves the functional form of $Q(\cdot)$. The most common candidate is the Cobb-Douglas framework. This functional form would yield an estimate of the output elasticity of labor that would be common to the set of firms to which the estimation pertains: $\hat{\theta}_{it}^L = \hat{\theta}^L$, hence, $\hat{\theta}_{it}^L = \hat{\theta}_{jt}^L$ for all firms $i$ and $j$, $i \neq j$, included in the estimation sample. It follows that the heterogeneity of firm markups and the ratio of the output elasticity of labor on its revenue share would simply reflect heterogeneity in the revenue share of labor: $\mu_{it}^L = \frac{\theta^X}{\alpha_{it}^L}$. Therefore, we prefer to use the translog production function because it generates markups whose distribution is not solely determined by heterogeneity in the revenue share of labor, as will be clear below.

To obtain consistent estimates of the output elasticity of labor $\theta_{it}^L$, we restrict our attention to production functions with a scalar Hicks-neutral productivity term and with technology parameters that are common across firms. Thus, we have the following expression for the production function:

$$Q_{it} = F(L_{it}, K_{it}; \mathbf{B}), \tag{6}$$

where $\mathbf{B}$ is a set of technology parameters to be estimated. Let $q_{it}$ be the translog production function:

$$q_{it} = \beta_l l_{it} + \beta_k k_{it} + \beta_{lk} l_{it} k_{it} + \beta_{kk} k^2 + \beta_{ll} l^2 + \omega_{it} + \varepsilon_{it}, \tag{7}$$

where smaller cases indicate the log transform, $\omega$ is a measure of the true productivity, and $\epsilon$ is true noise [5]. The estimation of vector $\mathbf{B}$ is challenged by the correlation of variable input $L$ with the productivity term $\omega_{it}$, which is known by the entrepreneur but not by the econometrician. The resulting endogeneity of labor $L$ would yield inconsistent estimates for vector $\mathbf{B}$. To overcome the problem of endogeneity, we use a control function approach, as in Olley & Pakes (1996) or Levinsohn & Petrin (2003), using demand for material to proxy for productivity:

$$m_{it} = m_t(k_{it}, \omega_{it}, ED_{it}), \tag{8}$$

[4]Although highly relevant, this issue represents a research program of its own; to our knowledge, all methods that measure market power need to assume a perfect market in at least one of the variable inputs.

[5]Note that we recover the Cobb Douglas (CD) production function in logs when omitting higher-order terms ($\beta_{kk}k^2, \beta_{ll}l^2$) and the interaction term $\beta_{lk}l_{it}k_{it}$.

where $ED_{it}$ is an indicator variable set to unity if the firm records positive exports and 0 otherwise. The inclusion of the firm's exporting behavior in the first stage of the estimation algorithm illustrates the need to control for firm characteristics that may affect factor demand (De Loecker & Warzynski 2012). Firm export behavior is an important feature that may affect both the quantity and the quality of the optimally desired input.

As in Olley & Pakes (1996), one can then invert the function and write productivity $\omega$ as follows:

$$\omega_{it} = h_t(k_{it}, m_{it}, ED_{it}). \tag{9}$$

Step 1 starts with the following estimation:

$$q_{it} = \phi(l_{it}, k_{it}, m_{it}, ED_{it}) + \varepsilon_{it}, \tag{10}$$

where the function form of $\phi$ is set to the third-order polynomials and a full vector of interactions. This first step is used to generate an estimate of expected output $\hat{\phi}_{it}$ and $\varepsilon_{it}$. Let productivity $\omega$ be the residual defined as follows:

$$\omega_{it}(\beta) = \hat{\phi}_{it} - \hat{\beta}_l l_{it} - \hat{\beta}_k k_{it} - \hat{\beta_{lk}} l_{it} k_{it} - \hat{\beta}_{ll} l_{it}^2 - \hat{\beta_{kk}} k_{it}^2. \tag{11}$$

Defining the law of motion for productivity as a first-order Markov process allows us to recover true innovation $\xi_{it}$ in the productivity equation:

$$\omega_{it} = g_t(\omega_{it-1}) + ED_{it} + \xi_{it}. \tag{12}$$

As suggested by Ackerberg et al. (2006), we use the following moments to obtain our estimates of the production function:

$$E\left(\xi_{it}\left(\beta\right)\left(\begin{array}{c} l_{it-1} \\ k_{it} \\ l_{it-1}^2 \\ k_{it}^2 \\ l_{it-1}k_{it} \end{array}\right)\right) = 0. \tag{13}$$

Equation (13) provides the orthogonality conditions to estimate (7) and obtain the estimated output elasticity of labor $\hat{\theta}_{it}^L$:

$$\hat{\theta}_{it}^L = \hat{\beta}_l + 2\hat{\beta}_{ll} l_{it} + \hat{\beta_{lk}} k_{it}. \tag{14}$$

From Equation 14, it is clear that $\hat{\theta}^L$ is firm specific because both $l$ and $k$ are themselves firm specific and reflect the production technology of the firm. We compute the revenue share of labor as follows:

$$\alpha_{it}^L = \frac{w_{it} L_{it}}{P_{it}\tilde{Q}_{it}}, \tag{15}$$

where $w$ represents the firm's average wage and $\tilde{Q} = \exp(\hat{\varepsilon})^{-1}Q$. Finally, with information on $\alpha_{it}^L$ and $\hat{\theta}_{it}^L$, one can compute the markups for each firm $i$ at time $t$:

$$\hat{\mu}_{it} = \frac{\hat{\theta}_{it}^L}{\alpha_{it}^L}. \tag{16}$$

### 2.2 Preliminary Evidence on Markups, Productivity, Export Behavior and the Euro in French Manufacturing

We use a panel database of French firms covering the 1994-2007 period. Data come from the annual survey of companies (EAE) led by the statistical department of the French Ministry of Industry on all manufacturing sectors. The survey covers all French firms with at least 20 employees in the manufacturing sectors. EAE data provide information on the financial income and balance sheet, from which we retrieve data on sales (corrected for stock variations), value added, labor, number of hours worked, capital stock, and materials. Information on the export behavior of firms in terms of destination countries is retrieved from the French custom database. Merging financial statements from EAE with the custom data yields a final database of approximately 235,000 firm-year observations. Appendix A provides more information on the data and the series of deflators used in this paper.

Table 1 presents the main variables (in logs) used in the computation of firm-year markups and productivity indices for all manufacturing sectors and by industry.[6] The use of the lagged value of labor as instruments trims the number of observations to 215,049. The average markup across all industries and over the time period is 33%, a value that is considerably higher than the one found in Bellone et al. (2009), which amounts to 12% between 1990 and 2004.[7] The computed markups are of similar magnitude as the average of 28% provided by De Loecker & Warzynski (2012, 28%) for the case of Croatian companies and compare well with the operating margin rate of 32.6% provided by INSEE, the French Office for National Statistics.[8]

[Table 1 about here.]

The overall computed means conceal substantial cross-industry heterogeneity. For example, the average markup in "Manufacture of Coke and Fuel" is higher than that found in "Printing and Publishing" by an order of 5 (60.5% and 11.4%, respectively). The F-statistics lead us to conclude

[6] All translog coefficients have been estimated by industry.

[7] The main explanation for this gap lies in the methodology used. Roeger's approach (Roeger 1995) uses variables that are expressed in first differences, a procedure known to produce a downward bias in the estimated parameters (the markup, in this case).

[8] The operating margin rate is measured at the firm level as the ratio of operating income over value added.

that markups, together with all firm-level variables, are industry specific.[9] Therefore, in the following, we transform all variables by subtracting the industry average. However, the industry-specific effect explains only 10% of the variance of firm-level markups. This leaves considerable room for firm-level explanatory variables that must explain the variance of markups. We mainly think of productivity, size and exposure to international trade.

The estimation of firm-year markups and TFP allows us to plot their dynamics over the whole time frame. The solid line of Figure 1 displays the mean, the markers "x" show the median, and the vertical lines show the interquartile range for each year. The substantial decrease in markups begins in 1999 and continues to 2003, remaining at this level until 2005. The years 2006 and 2007 witness a rise in the average markups. Although surprising, this rise is fully consistent with the evolution of the operating margin rate as provided by INSEE.[10]. It is difficult to conclude that there is a significant change in the interquartile range. Hence, whether the introduction of the euro is associated with changes in the distribution of markups across firms remains an open issue.

[Figure 1 about here.]

**The introduction of the euro** may have had two opposite and sequential effects on the level of markups. First, the launch of the euro in 1999 is similar to reducing transaction costs for firms that trade with other euro countries. For these firms, the new exchange rate fixity is similar to a cancellation of all menu costs associated with changes in currency. Hence, one can expect a positive effect of the establishment of the euro on markups for eurozone exporters, the extent of which depends on whether firms choose to pass through the cost reduction to prices or to increase their markups. To capture this opportunity effect, we set variable $d_{p99}$ to unity if year $t$ exceeds 1999 and 0 otherwise: $d_{p99} = 1$ if $t \geq 1999$ and $d_{p99} = 0$ if $t < 1999$. Second, in 2002, the introduction of the euro to all economic agents – including final consumers – must be considered yet another step toward a more integrated European market, leading to an increase in product market competition. Because consumers became more able to arbitrate between products from different euro countries, it enlarged the European market, causing an increase in competition from foreign firms. This competitive effect is reinforced by the fact that the creation of the eurozone increased the profit opportunities associated with access to a larger market and acted as an incentive for new firms to enter the market. Although we expect the positive effect of the introduction of the euro on markups to begin in 1999, we expect the negative

[9] Although statistically significant, there is considerably less inter-industry variance in the case of TFP ($\omega$).

[10] See `http://www.insee.fr/fr/themes/tableau.asp?reg_id=0&id=180`.

effect to occur after 2002. Hence, we define variable $d_{p02}$ as unity if the year $t$ exceeds 2002 and 0 otherwise.

**The export status** is provided by the French customs database, which allows us to recover information on the volume of exports by destination. Table 2 provides descriptive statistics about the participation rates, the export intensity, and the entry and exit rates by destination markets and time period[11]. Over the 1995-2007 period, 58% of companies exported to destination markets located in the eurozone ($Dest_{EZ}$) and 68% exported to markets outside the eurozone ($Dest_{NEZ}$). These figures amount to an export participation rate of almost 73%. The overall export intensity reaches 28% and is almost equally divided between eurozone and non-eurozone markets. Moreover, because the entry rate exceeds the exit rate across all destination markets, this time period witnessed the net entry of firms into export markets and an increase in the export participation rate.

Producing these figures by time period leads us to observe that the two time periods of 1999-2001 and 2002-2007 conceal contrasted statistics. During 1999-2001, the sudden rise in export participation to almost 74% is mainly due to the net entry of firms into eurozone markets at the expense of non-eurozone markets. This rise is accompanied by an impressive increase in the export intensity of firms supplying eurozone markets, from 12.7% to 15%, and its concomitant decrease from 15.7% to 14.4% for non-eurozone markets. This preliminary evidence is consistent with the idea that the launch of the euro in 1999 was tantamount to reducing export costs to eurozone markets, thereby favoring the entry of firms. In the last period of 2002-2007, both the participating rate and the export intensity of firms exporting to the eurozone decreased, indicating the net exit of firms from eurozone markets. This finding is consistent with the idea that the introduction of a fiduciary euro in 2002 increased competition across all eurozone markets and led to the net exit of the least competitive firms.

[Table 2 about here.]

Using information about the destinations of exports, we define three exclusive dummy variables for the export status of firms. We distinguish non-exporters from exporters to the eurozone exclusively ($d_{EZ}$), exporters outside the eurozone ($d_{NEZ}$), and exporters to both the eurozone and outside the eurozone, which we call global exporters ($d_G$). Eurozone exporters are those firms that export in the eurozone, whereas global exporters have at least one destination outside the eurozone. Non-eurozone exporters have customers exclusively outside the eurozone. Global exporters, eurozone exporters, and non-eurozone exporters represent 53%, 5% and 15% of French manufacturing firms, respectively.

[11] See the table footnote for information on the computation of the series of ratios

Table 3 displays summary statistics on computed markups per period and per export status. We observe that both the markup and the TFP of exporters exceeds that of non-exporters significantly, a stylized fact that is now well documented. Exporters self-select into export markets because their productivity advantage provides financial resources necessary for entry into foreign markets (ISGEP 2008). Exporters also charge higher markups, either as a result of their productivity advantage or as an expression of a price elasticity that differs in international markets (De Loecker & Warzynski 2012). Moreover, all means regarding firm markups and TFP indices differ significantly across export status.

[Table 3 about here.]

The top panel of Figure 2 displays the kernel distributions for markups by time period. We observe a leftward move, identifying a decrease in markup without a change in variance among firms. Table 4 provides the results from the Student t-tests and Kolmogorov Smirnov (KS) tests addressing the questions of homogeneity across export status and across years. The former tests address the issue of equality of group means, whereas the KS test addresses the issue of stochastic dominance of one distribution over another. All tests are significant. The KS tests show that the distribution of the first period (1995-1998) stochastically dominates that of the second period (1999-2001) and that the distribution of the second period stochastically dominates that of the last period (2002-2007). Hence, most firms have experienced a decline in markups over time.

[Figure 2 about here.]

Similarly, the bottom panel of Figure 2 shows kernel distributions for markups by export status. We observe a rightward move from non-exporters to global exporters. The KS tests show that the distribution of global exporters stochastically dominates that of exporters to the eurozone. The distribution of EZ exports dominates that of NEZ, which dominates that of non exporters. Significant differences in markups across export status are difficult to interpret. Bellone et al. (2014) show that the size of the destination market, its wealth and its geographic distance to the local market can affect the level of markups.

[Table 4 about here.]

Table 3 also provides the mean values of both markups and productivity by export status and by time period. At first glance, decreases in average markup and increases in productivity from the first to the third period are observed. This finding is consistent with the idea that productivity increased with time and that the establishment of the common market and the launch of the euro led to an increase in competition and, hence, a decrease in markups. These trends are observed irrespective of the export status for markups or for the productivity indices. Nevertheless, it remains difficult to draw conclusions regarding the heterogeneity in markups: interquartile ranges (IQR) in markups seem to have increased from 1995-1998 to 1999-2001 and then decreased in the last period of 2002-2007. Heterogeneity in TFP, both in terms of variance and IQR, increased over time. These trends are observed irrespective of the export status.

These findings provide preliminary evidence that the introduction of the euro is concomitant with significant variations in markups, which may differ across firms. The next section intends to provide econometric evidence of the differentiated impact of the establishment of the common currency on firm markups.

# 3 Markups and Firm-level Responses to the Introduction of the Euro

## 3.1 Econometric Specification

The empirical model explains markups $\mu$ of firm $i$ at time $t$, as derived from Equation 16, conditional on firm export status $d_{\mathbf{X}}$ and the time period of observation $d_{\mathbf{p}}$. The empirical model reads

$$\begin{aligned}\mu_{it} = &\sum_{\mathbf{X}} \beta_{\mathbf{X}} d_{X,it} + \sum_{\mathbf{p}} \beta_{\mathbf{p}} d_{\mathbf{p,it}} + \sum_{\mathbf{X,p}} \beta_{\mathbf{X,p}} (d_{\mathbf{X}} \times d_{\mathbf{p}})_{it} \\ &+ \mu_0 + \mu_1 \omega_{it} + \lambda t + \mathbf{BC} + \mathbf{u_i} + \epsilon_{\mathbf{it}},\end{aligned} \tag{17}$$

where $\mathbf{X}$ = {EZ;NEZ;G}, $\mathbf{p}$ = {p99;p02}, letter $d$ denotes binary variables, $\mu_0$ is the constant, $\omega$ is firm TFP and $\mu_1$ its associated parameter, $t$ is a linear time trend and its parameter $\lambda$ provides the percentage-point change in $\mu$, $\mathbf{C}$ is a vector of additional controls, and $\mathbf{B}$ is its associated vector of parameter estimates. The term $u_i$ is a firm fixed effect controlling for unobserved heterogeneity, and $\epsilon_{it}$ is the error term. Because the two time-period dummy variables $d_{p99}$ and $d_{p02}$ overlap, interpretation of their associated parameters will be additive for all years greater than 2002. These two dummies also control for the overall impact of the introduction of the euro on markups, irrespective of the export status. The parameters of interest are the series of $\beta$s, whose value indicates the response of firms according to their export status and the period of interest.

This model is similar – but not identical – to the so-called difference-in-differences model (DID, henceforth), which classically compares a treated group with a control group. However, it is not clear what defines the treatment and which firms constitute the treated and the control groups. If the euro is the treatment, the introduction of the euro applies to all companies regardless of their export status, whereas in a DID model, the treatment applies to the treated group only. If eurozone exporters are the treated group, it is not clear what the counterfactual group should be. Although non-exporting firms are obviously eligible, it is not clear how firms that export exclusively outside the eurozone should be considered. Therefore, specification 17 *resembles* a DID model but is not a model in its most standard fashion.

Our intent to isolate the effect of the establishment of the euro conditional on export destination leads us to interact the euro period with the three exclusive export status. We expect $\beta_{\mathrm{EZ,p99}}$ to be positive following the decrease in transaction costs (opportunity effect), whereas $\beta_{\mathrm{EZ,p02}}$ should be negative (competitive effect). Heterogeneity in firm responses to the shock induced by the introduction of the euro is provided by first comparing $\beta_{\mathrm{EZ,p99}}$ with $\beta_{\mathrm{NEZ,p99}}$ and $\beta_{\mathrm{G,p99}}$ to identify the opportunity effect and second comparing $\beta_{\mathrm{EZ,p02}}$ with $\beta_{\mathrm{NEZ,p02}}$ and $\beta_{\mathrm{G,\ p02}}$ to identify the differential impact of the competitive effect with respect to export status.

Bertrand et al. (2004) warn about the use of a DID-type model with panel data because the outcome variable may be serially correlated. In such a case, the resulting standard errors for the coefficients of interest are known to be inconsistent. Their proposed solution is twofold. First, one can estimate corrected standard errors by bootstrapping samples on firms and keeping the entire series of information for firm $i$. This approach preserves the autocorrelation structure, which is then corrected by clustering the standard errors by firms. Second, the alternative solution is to average the data by time period to minimize the time-series information. Because the resulting sample consists of three time periods, the specification with a firm fixed effect still holds.

[Table 5 about here.]

[Table 6 about here.]

Vector **C** includes variables that may presumably affect the level of markups above and beyond their export destinations and volumes. We include firm size measured by overall sales , the share of imports of an industry whose provenance is from low-wage countries, and the degree of openness of a given industry, defined as the share of trade – imports and exports – over the value added of that industry. Specifications 17 can be estimated by least squares with standard errors clustered by firms.[12] The definition of variables

[12]We follow Bernard et al. (2006), who classify low-wage countries as countries for which per capita GDP does not exceed 5% of that of the US. The list includes Afghanistan, Albania, Angola, Armenia, Azerbaijan, Bangladesh, Benin, Bhutan, Burkina Faso, Burundi, Cambodia, Central African Republic, Chad, China, Comoros, Congo, Equatorial, Eritrea,

and the data sources are provided in Table 5, and the descriptive statistics are provided in Table 6.

### 3.2 Export Status and the Introduction of the Euro

We first proceed with least-squares regression by sequentially introducing the key variables. We do not introduce any particular control apart from TFP, so these results should be seen as merely descriptive statistics controlling for TFP. The results are displayed in Table 7.

In Column (1), the base markup for all firms (the constant) amounts to 1.107, which implies a pooled price-cost margin of 10.7%. We observe a positive correlation between $\mu$ and $\omega$ (+0.143), which suggests that more productive firms enjoy higher markup. Exporters display a 0.118 markup premium.[13]. These preliminary remarks are coherent with the empirical literature, which repeatedly shows that exporters outperform non-exporters in several dimensions, such as productivity, wage per employee, size and markups.

Regarding the introduction of the euro, the two 'euro dummies' are negative and significant, which suggests that the establishment of the common currency acted as a competitive shock and forced firms to decrease their markups. Parameter $\hat{\beta}_{p99}$ is negative, suggesting that the competitive effect is effective prior to 2002. This could be caused by firms anticipating the 2002 eurozone final establishment from both incumbents and new entrants, with the former adjusting their markups downward to prepare the run-up phase toward a more integrated market and the latter entering the eurozone markets before the introduction of the euro due to decreased sunk entry costs in the eurozone markets.

Observe that parameter $\hat{\beta}_{p02}$ is four times as large as $\hat{\beta}_{p99}$. This difference in magnitude is persistent across specifications (1) to (4). This finding suggests that on the supply side, the competitive effect is stronger in 2002, whereas on the demand side, the ability of consumers from 2002 onwards to compare similar products across eurozone markets had a substantial impact. The figures from Table 3 indicate that firm entry into the eurozone market was more important in the first period (1999-2001) than in the second (2002-2007). This finding suggests an over-adjustment effect of numerous firm entries followed by various firm exits in the second period due to the strengthening of the competition.

In Column (2), the export status distinguishes EZ, NEZ and global ex-

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Ethiopia, Gambia, Georgia, Ghana, Guinea, Guinea, Guinea-Bissau, Guyana, Haiti, India, Kenya, Lao PDR, Leone, Lesotho, Madagascar, Malawi, Maldives, Mali, Mauritania, Moldova, Mozambique, Nepal, Niger, Pakistan, Rwanda, Samoa, Sao, Sierra, Somalia, Sri Lanka, St. Vincent, Sudan, Togo, Tome, Uganda, Vietnam, and Yemen.

[13]A firm becoming an exporter and having an average level of productivity should have a markup equal to $\mu = 1.107 + \bar{\omega} \times 0.143 + 0.118$ before the introduction of the euro

porters. All types of exporters enjoy higher markups, with global exporters enjoying higher markups and exporters outside the eurozone having lower markups than EZ exporters. Interpretation is not straightforward because information on the destination market is likely to capture two opposite effects (Bellone et al. 2014): one positive (the wealth of the destination country, which acts as a proxy for product quality) and one negative (the distance to the destination market, which decreases the markup by increasing distribution costs).

In Columns (3) and (4), we explore the contrasted effect of the establishment of the euro on markups by interacting the period dummies with the export status. In Column (3), the interaction term concerns only the first period of 1999-2007. This interaction is positive and significant for EZ exporters (+0.016) and negative and significant for global exporters (−0.010). Summing the parameter estimates with the autonomous effect $\hat{\beta}_{p99}$, we observe that the establishment of the euro is associated with a decrease in markups of 1.1 percentage points for EZ exporters and 3.7 percentage points for global exporters.

Interacting the export status with the second period of 2002-2007 (Column 4) eliminates the significance for all interacted effects in the first period but leads to similar results for the second period. This interaction is positive and near significant for EZ exporters (+0.015) and negative and significant for global exporters (−0.028). Summing the parameter estimates with the autonomous effect $\hat{\beta}_{p02}$, we observe that the general use of the euro in 2002 is associated with a severe decrease in markups of 9.4 percentage points for EZ exporters and 13.7 percentage points for global exporters. This finding suggests that the euro shock was milder for exporters to the eurozone than other firms and more severe for global exporters.

[Table 7 about here.]

The estimation of equation 17 represents the key results of the specification. These results are displayed in table 8. In Column (5), in addition to a linear time trend, we introduce a series of firm-level and industry-level variables, and we control for unobserved firm heterogeneity. First, we observe that the 2002 shock is three times as large as the 1999 shock. Interactions with the export status reveal that the 1999 shock allowed eurozone exporters to partially compensate the 3.3 percentage point decrease in markup by 2.2 points so that the net effect amounts to a decrease of 1.1 percentage points.[14] Our interpretation is that the 1999 shock represented an opportunity for eurozone exporters to rejuvenate their margins because the newly fixed exchange rate was tantamount to reducing transaction costs, which arguably, represent a significant part of their export costs.

[14]For clarity, we do not report the autonomous effect of the export status variables. In a firm fixed-effect model, the estimated parameter estimates only report the effect of a change in the export status, whereas we are interested in the interaction effect.

Although global exporters also experienced a reduction in export fixed costs proportional to their relative involvement in the eurozone, their markups did not increase in 1999 relative to other firms. Instead, the 2002 shock was more severe by $-2.4$ percentage points, with the net effect reaching $-12.5$ percentage points. Our understanding is that after 2002, the appreciation of the euro was detrimental to their margins. There is an issue with regard to non-eurozone exporters: why would they not undergo the same effect as global exporters, although they too experienced a similar change in the euro? Our tentative explanation is as follows: global exporters export differentiated products (similar to EZ exporters), whereas NEZ exporters are likely to sell products of a lower quality that are more substitutable and, consequently, have lower markups. Indeed, descriptive statistics show that for export status, EZ exporters have the lowest markups. Because they have low price-cost margins, their ability to adjust by decreasing their markups is mechanically reduced. Instead, firms are forced to withdraw from export markets. This hypothesis is supported by the descriptive statistics in Table 3, which display a clear decrease in the share of outside eurozone exporters over the period, from 16.8% in the first period to 14.8% in the third period. In the remaining regressions, we control for entry into and exit from domestic and export markets to account for any potential selection effect on the estimated coefficients.

The sign of the control variables conforms to our intuition. Firm size is positively associated with markups, suggesting that the exploitation of economies of scale allows the firm to price above its marginal cost of production. Firm size is similar to controlling for a host of firm-level unobserved characteristics, notably monopsony power in the input factors markets, which may affect firms' product market power. However, including size does not affect our previous conclusion on the euro effect. Efficiency in production (TFP) allows firms to increase their margins. Finally, we observe that competition from low-wage countries has a negative and significant on the determination of markups. This finding indicates that the entry of new players in the market arena forced firms to discipline themselves in their price behavior. Sector openness and grasping the globalization of markets and the international division of labor are not related to price-cost margins.

[Table 8 about here.]

Columns (6) to (9) display the results of the same specification using different sub-samples to account for any change in markups due to selection effects. Column (6) controls for industry churning by excluding firms that entered or exited domestic markets between 1998 and 2003. The results are fully consistent with prior estimates, inflating the estimated opportunity effect of 1999 for EZ exporters by a factor of 1.5 (+0.031) with respect to Column (5). Column (7) represents the key result of the table, where only firms that do not change export status between 1998 and 2003 are included in the sample.[15] Following the recommendation of Bertrand et al. (2004), we bootstrap on individual firms to retain all the time-series information by randomly selected cross-section.[16] First, we observe that the 2002 shock is double that of 1999. The true competitive shock occurred in 2002. Adding the two effects amounts to a substantial 13.8 percentage point decrease in price cost margins.

Although less efficient, parameter $\hat{\beta}_{EZ,p99}$ (+0.052) indicates that eurozone exporters could entirely compensate for the competitive effect of the euro on markup ($\hat{\beta}_{p99} = -0.046$). Observe that the interactions with other export status suggest that they too could accommodate the competitive shock to a milder threshold ($\hat{\beta}_{NEZ,p99} = +0.014$ and $\hat{\beta}_{G,p99} = +0.016$). What these two groups have in common relative to EZ exporters is that they may have benefited from the 1999-2002 depreciation of the euro (notably, with respect to the US dollar), leading them to increase their markups when keeping local-currency prices constant. The 2002 shock was more severe ($\hat{\beta}_{p02} = -0.092$) for all firms in general and for global exporters in particular ($\hat{\beta}_{G,p02} = -0.028$), leading to a decrease in their markups amounting to 12 percentage points.

Another strategy is to minimize the time-series information and compute averages by period and by individual firms (Bertrand et al. 2004). Because of the three time periods, the panel dimension is 3, and we compute robust standard errors to account for the residual serial correlation in the error term. The results displayed in Column (8) corroborate the previous remarks. Finally, in Column (9), we produce a sample of firms that matches the characteristics of eurozone exporters in TFP, size, competition from low-wage countries and openness to international trade in 1997, trimming the number of observations to 8,000, which corresponds to 790 companies.[17] The objective here is to produce a counterfactual sample that shares the characteristics of EZ exporters in various dimensions (with the exception of markups $\mu$) to interpret differences in export status as heterogeneous firm responses to a common shock. Although this exercise produces similar results qualitatively, some interaction variables are amplified in magnitude ($\hat{\beta}_{EZ,p99} = +0.090$, $\hat{\beta}_{NEZ,p99} = +0.034$ and $\hat{\beta}_{EZ,p02} = -0.138$). When comparing EZ exporters with firms of similar characteristics, we observe the largest variations in markups following the introduction of the euro concerns eurozone exporters.

Taking stock of our results thus far, we conclude that the two shocks related to the establishment of the euro substantially increased competition in French manufacturing industries, implying a dramatic decrease in markups

[15] We classify firms according to their export behavior between 1998 and 2003 irrespective of their export status for the remaining years.

[16] In the terminology of the authors, we "block bootstrap".

[17] See Appendix B reporting the results from the matching procedure.

of *circa* 14 percentage points. The 2002 shock was significantly more severe than the initial one. Importantly, both shocks were asymmetrical, benefitting eurozone exporters in 1999 relative to all other companies and harming global exporters more severely after 2002. It is consistent with our intuition that the introduction of the euro in 1999 significantly reduced export costs within the eurozone and was an opportunity for such exporters to increase their markups. The control variables behave as expected. There is an issue with the variable openness, suggesting that globalization may have discriminated firms in some fashion.

### 3.3 Exposure to the Euro Shocks

The previous econometric specification distinguishes firms that export to the eurozone ($d_{EZ}$) from those that export outside the eurozone ($d_{NEZ}$) and those that export globally ($d_G$). One could object that vector $\beta_{\mathbf{X},\mathbf{p}}$ fails to accurately grasp the response of firms to the euro shock. First, global exporters also export to the eurozone markets; therefore, both groups $d_{EZ}$ and $d_G$ are potentially affected by the introduction of the euro, whereas only outside eurozone exporters directly experienced the change in the euro value. By using the share of export to the eurozone, we can focus on the eurozone competitive effect only and discard the effect of the appreciation of the euro. Second, the dichotomous nature of dummy variables makes no use of the amount of exports to the eurozone. This information acts as an exposure to the shock induced by the introduction of the euro, allowing the markup to adapt gradually with the share of exports to the eurozone. Therefore, an alternative empirical model could read as follows:

$$
\begin{aligned}
\mu_{it} =& \beta_1 \alpha_{EZ,it} + \sum_{\mathbf{p}} \beta_{\mathbf{p}} d_{\mathbf{p},\mathbf{it}} + \sum_{\mathbf{p}} \beta_{\mathbf{p},\mathbf{EZ}} \big(\alpha_{EZ} \times d_{\mathbf{p}}\big)_{it} \\
&+ \beta_2 \ln X_{it} + \mu_0 + \mu_1 \omega_{it} + \lambda t + \mathbf{BC} + \mathbf{u_i} + \epsilon_{\mathbf{it}},
\end{aligned}
\tag{18}
$$

where letter $d$ and vector $\mathbf{p}$ are defined as previously, $\ln X$ is the log-transformed value of overall exports[18], and $\alpha_{EZ}$ is the share of exports to the eurozone: $\alpha_{EZ} = X_{EZ}/X$. Variable $\alpha_{EZ}$ is interpreted as the exposure of the firm exporting to the eurozone to the introduction of the euro, and $\beta_{EZ,p99}$ and $\beta_{EZ,p02}$ are the responses by firms from a unit variation in $\alpha$. Table 9 displays the results for the same sub-samples defined in the previous subsection. In our comments, we concentrate on Column (12), including firms that are stable in their export behavior and excluding firms that entered or exited the market between 1998 and 2003. The remaining columns are displayed for robustness checks.

[18] To account for firms that do not export where $X = 0$, the log transform is performed as follows: $\ln(X + 1)$.

[Table 9 about here.]

The two dummy variables for the introduction of the euro keep their negative sign and significance; the pro-competitive effect of the post-2002 era is far more dramatic for price-cost margins. We observe that firms that export to the eurozone have higher markups and that the larger the share of exports to the eurozone, the higher the markups. The interaction of $\alpha_{EZ}$ with the two time period dummies yields consistent results with our previous findings. The 1999-2001 period, which established fixed exchange rates across the currencies of the eurozone, was tantamount to reducing export costs to the eurozone and providing an opportunity for such exporters to increase their margins. Conversely, the introduction of the currency in general use in 2002 acted as a very strong pro-competitive effect for intensive exporters to the eurozone. This finding sheds light on the result of Table 8: global exporters suffer the most from the competitive effect, not only because they have to face the appreciation of the euro but also because global exporters are those with the largest export intensity to the eurozone on average.[19]

Another satisfaction comes from the stability of the parameter estimates for the remaining control variables, such as TFP, size, low-wage imports and the absence of significance of openness. This finding reinforces our confidence in our results and in the idea that the introduction of the euro had a strong pro-competitive effect on markups but that this effect was asymmetrical, depending on the export behavior of firms and the degree of exposure to the euro shock.

### 3.4 Heterogeneity in Firm Response to the Introduction of the Euro

Quantile regression techniques are an alternative to least squares that provide an estimation of the heterogeneous impact of any explanatory variable on the dependent variable. Our motivation comes from the fact that the introduction of the euro may be heterogenous, not only in terms of exposure to exports but also in its impact on markups depending on the initial market power of firms. In essence, the spirit of quantile regression is to produce quantile-specific marginal effects $\hat{\beta}_\tau$, where $\tau$ is the desired quantile.

To account for unobserved heterogeneity in a quantile regression setting, we rely on the two-step estimator developed by Canay (2011). In the first step, we perform a least squares regression in which we introduce firm fixed

[19]Unreported statistics show that the ratio of exports to the eurozone in sales is 11.7% and 13.6% for EZ and global exporters, respectively. The contrast is sharper when considering the median of the ratio, with 2.3% and 7.3% for EZ and global exporters, respectively. Because global exporters are considerably larger than EZ exporters, the former outperform the latter in the volume of exports by a factor of 10.

effects and then define $\breve{\mu}_{it} = \mu_{it} - \hat{u}_i$, where $\hat{u}_i$ is the least square estimate of the firm fixed effect. In the second step, we estimate a traditional quantile regression on the transformed dependent variable $\breve{\mu}_{it}$ as follows:

$$\begin{aligned} Q_{\breve{\mu}}(\tau|\mathbf{Z}) &= \beta_1(\tau)\alpha_{EZ} + \sum_{\mathbf{p}} \beta_{\mathbf{p}}(\tau) d_{\mathbf{p}} + \sum_{\mathbf{p}} \beta_{\mathbf{p},\mathbf{EZ}}(\tau)\big(\alpha_{EZ} \times d_{\mathbf{p}}\big) \\ &+ \beta_2(\tau) \ln X + \mu_0(\tau) + \mu_1(\tau)\omega + \lambda(\tau) t + \mathbf{B}(\tau)\mathbf{C} + \mathbf{F}_{\epsilon}^{-\mathbf{1}}(\tau), \end{aligned} \tag{19}$$

where $\mathbf{Z}$ is the vector of explanatory variables, subscripts $i$ and $t$ have been omitted for the sake of clarity, $\tau \in (0,1)$ and $F_{\epsilon}^{-1}(\tau)$ denotes the common distribution function of the errors. In his contribution, Canay (2011) shows that estimation of the covariance matrix can be achieved by bootstrapping on the set of observations. Table 10 displays the estimations of quantiles (Columns 15 to 19) for various values of $\tau$ and recalls the results from the least squares solution (Column 12 from Table 9).

[Table 10 about here.]

We use the restricted sample, including firms present at least from 1998 to 2003 that do not change export status. The quantile estimates indicate that the introduction of the euro has been far more dramatic for firms with an initially high market power, decreasing their markups by almost 6.7 percentage points for the $90^{th}$ percentile of the distribution in the 1999-2001 period and by 11.2 percentage points in the post-2002 era, leading to an overall effect of an almost 18 percentage point decrease for firms located in the right tail of the distribution. Conversely, firms with a lower market power have decreased theirs by less than 9.1 percentage points since 1999 ($\hat{\beta}_{p99}(\tau = 0.1) + \hat{\beta}_{p02}(\tau = 0.1) = -0.091$). Therefore, the introduction of the euro and the subsequent increase in competition have had an important heterogeneity-decreasing effect on the distribution of markups.

A more thorough examination of the heterogeneity of the overall euro effect is found in Figure 3. In the upper and bottom left quadrants, the actual effect is displayed for each of the $5^{th}$ to the $95^{th}$ percentiles. The grey area displays its corresponding $95^{th}$ confidence interval. The horizontal solid and dashed lines represent the OLS effect with a $95^{th}$ confidence interval. We observe that the actual OLS effect conceals a great deal of heterogeneity in the euro effect, where the most severe downward adjustments in the markups are to be found in the largest markups. This must be expected; theoretically, firm markups have a lower bound at unity. Thus, the magnitude of adjustment for firms with initially lower markups is narrower, whereas firms with higher markups can trim their margins more extensively.

Importantly, the sign of the slope depicted by the graph across the quantiles is another indication of the effect of a given variable on the variance of markups. More precisely, a negative slope is tantamount to observing an effect that decreases the variance of the dependent variable because observations located in the right tail of the distribution adapt at a lower pace than observations located in the left tail. In the case of markups by boosting competition, the introduction of the euro has clearly reduced heterogeneity in market power across firms. The positive 1999 effect of exposure and the negative 2002 effect of exposure are homogenous across firms.

[Figure 3 about here.]

Firm characteristics, whether size or productivity, have a heterogeneity-increasing impact on markups. The quantile regression tells us that a change in size has a larger impact on markups when firms have an initially strong market power – a large markup. Although statistically significant, the change in the marginal effect from the $1^{st}$ to the $9^{th}$ decile is economically small. A similar heterogeneity-enhancing effect is found for productivity $\omega$. Computing the difference $\hat{\beta}_\omega(\tau = 0.1) - \hat{\beta}_\omega(\tau = 0.9) = 0.179$, we conclude that productivity gains are far more beneficial for firms with initially higher markups. Figure 4 illustrates clearly and neatly the marginal effect of size and of productivity and the increasing slope, indicating that both size and productivity increase heterogeneity in markups across firms.

[Figure 4 about here.]

The positive slope across quantiles of the marginal effects of the variables on international competition (LW import share) and globalization (openness) provides evidence of their heterogeneity-enhancing effect on markups. First, looking at the effect of competition from low-wage countries, we observe that firms with lower markups have suffered the most from the emergence of such countries. If one reads markup as an inverse proxy for the degree of substitution, this can be interpreted as a sign that these countries started by producing goods where the substitution effect is highest. Computing $\Delta\hat{\beta}_{LWI}$ yields an impressive 24.5 percentage point differential. The upper right quadrant of Figure 4 illustrates this impressive differential. Finally, the sign of "openness" depends of the firm's location in the markup distribution. The effect is negative for firms with initially low markups and positive for firms with higher markups. The bottom right quadrant of Figure 4 illustrates the fact that the difference in sign is located at the median of the distribution. Our interpretation is that globalization represents a markup opportunity for firms with larger markups and has a pro-competitive effect only for firms with lower markups. Although not at the core of our paper, this result is crucial to understand the role of globalization in exacerbating differences in efficiency across firms.

## 4 Conclusion

We have investigated the effects of the introduction of the euro on the markups of French manufacturing firms. Relying on the recent methodology from De Loecker & Warzynski (2012), which allows us to compute time-varying, firm-specific markups, we show that the introduction of the euro has led to a significant decrease in the average markup of almost 14 percentage points. We find that the generalization of the use of the euro in 2002 had a more pronounced pro-competitive impact on markups (almost three times as large) as the mere establishment in 1999 of fixed exchange rates among eurozone countries.

Importantly, the launch of the euro has been an asymmetric shock for French manufacturing firms. First, depending on their export destination, the euro shock benefitted eurozone exporters in 1999 relative to all other companies and harmed global exporters more severely for the 2002 period. Our interpretation is that the introduction of the euro has had two opposite effects on the level of markups: (i) by reducing transaction costs for firms trading with euro-countries, the euro represented an opportunity for eurozone exporters to increase their price-cost margins relative to other firms in the sample; and (ii) by increasing competition, we find that all firms significantly reduced their markups.

The above conclusions hold regardless of the export status. However, global exporters are those whose markups suffered most from 2002 onward. This is because since global exporters are those who export most to the eurozone, they were the most exposed to the competitive shock of 2002. Moreover, the appreciation of the nominal exchange rate vis-a-vis other currencies represented a negative competitive shock for French products; they coped with this shock by compressing their margin. Non-eurozone exporters, who were not concerned with the opportunity effect or the competitive effect, did not experience a decrease in their markups relative to non-exporters. We believe that this illustrates a narrower window of adjustment to reduce their margins as a response to the appreciation of the euro, which forced some of them to exit foreign markets.

Second, quantile regressions show that the creation of the euro led to a reduction in the heterogeneity of firm markups. Because firm markups have a lower bound at unity, the magnitude of adjustment for firms with initially high markups is simply larger. As such, in addition to moving the whole distribution of markups leftward, the introduction of the euro induced a more uniform distribution of market power across firms. Increased competition from low-wage countries and openness are strong drivers of asymmetric position in markets and benefit firms that have products with lower degrees of substitution, that is, higher markups.

Our results are robust to a range of specifications, whether we consider the export status in a DID model or whether we quantify the exposure of the firm to the establishment of the euro, and across various sub-samples, controlling for firms exiting from or entering into either domestic markets or foreign destinations or considering non-eurozone exporters as a relevant counterfactual.

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# Appendix A. Data Appendix

All nominal output and inputs variables are available at the firm level. Industry level information is used for price indexes, number of hours worked and depreciation rates of capital.

**Output**. Our Output variable, $Q$, is value added deflated by sector-specific price indexes for value added. These indexes are available at the 2-digit level published by INSEE (French Office of Statistics).

**Labor**. We define our labor variable, $L$, as the number of effective workers multiplied by the the number of hours worked in a year. The annual series for worked hours are available at the 2-digit industry level and provided by *GGDC Groningen Growth Development Center*. This choice was made because there are no data on hours worked in the EAE datasets.

**Capital input** Capital stocks, $K$, are computed using information on investment and book value of tangible assets (we rely on book value reported at the end of the accounting exercise), following the traditional permanent inventory methodology:

$$K_t = (1 - \delta_{t-1})\, K_{t-1} + I_t \tag{A.1}$$

where $\delta_t$ is the depreciation rate and $I_t$ is real investment (deflated nominal investment). Both investment price indexes and depreciation rates are available at the 2-digit industrial classification from the INSEE data series.

**Intermediate inputs**. Intermediate inputs, $M$, are defined as purchases of materials and merchandise, transport and travel, and miscellaneous expenses. They are deflated using sectoral price indexes for intermediate inputs published by INSEE.

**Revenue shares**. To compute the revenue share of labor, we rely on the variable *wages and compensation*. This value includes total wages paid to salaries, plus social contribution and income tax withholding.

## Appendix B. The Matching Method

We choose the matching method to pre-process our sample in order to obtain a set of observations sharing similar characteristics. These characteristics are supposed to affect the response of firms in their markups above and beyond the export status. We mainly think of firm size ($Y$), all factors of production $K$, $L$ and $M$, and TFP $\omega$, being a non linear transformation of the series of output and inputs. We do not include a time fixed effect because we perform the matching procedure for year 1997, two years prior to the introduction of the euro.

We wish to produce a set of observations similar to firms exporting essentially in the eurozone, which we then consider as the treated group. All other firms are then considered non treated. Let $T$ denote the treated group, we set $T = 1$ if $\alpha_{EZ} > .9$, 0 otherwise, where $\alpha_{EZ}$ is the share of exports to the eurozone. This produces a dataset of 6,367 companies in 1997, 204 of which exports essentially or exclusively to the eurozone. The matching is then performed by:

$$Pr(T = 1|\omega, y, k, l, m) = \Phi\big(\omega, y, k, l, m\big) \tag{B.1}$$

where lower case denote the log-transformed values of the variables previously defined, and $\Phi$ is the logit function. The model is estimated by maximum likelihood and produces the propensity score, which is simply the predicted probability of a company being treated. We then compose the counterfactual by including the three nearest neighbors, chosen from the non treated group for each treated firm in our sample.

Table B1 presents the results of the matching procedure. The first column present the estimated coefficients from model B.1 on all 6,367 observations from year 1997. It shows that there exists important differences between treated and non treated firms of all dimensions, with the exception of labor. Pairwise Student t-tests exhibit also systematic differences across the two groups. Using results from the logit model to produce the propensity score yields a dataset composed of 790 firms, 586 of which do not export or export a minor part of their production to the eurozone. Performing the logistic regression on the treated and counterfactual set of firms and the pairwise student t-tests show no systematic differences between the two groups.

Table B1: Results from the Matching Procedure (♯ Treated = 204)

| | Logit (1) | $\bar{x}_{T=1}$ | $\bar{x}_{T=0}$ | Logit (2) | $\bar{x}_{CF}$ |
|---|---|---|---|---|---|
| $\omega$ | 1.367* | 1.424 | 1.395 | .951 | 1.416 |
| | (.779) | | [1.49] | ( 1.308) | [0.32] |
| $y$ | -1.342*** | 9.323 | 9.167 | -.309 | 9.405 |
| | (.488) | | [1.55] | (.584) | [-0.61] |
| $k$ | .474*** | 8.026 | 7.646*** | .023 | 8.117 |
| | (.102) | | [3.18] | (.129) | [-0.60] |
| $l$ | .469 | 11.93 | 11.88 | .428 | 11.99 |
| | (.411) | | [0.65] | (.589) | [-0.56] |
| $m$ | .594* | 8.825 | 8.600** | -.029 | 8.908 |
| | (.334) | | [1.96] | (.453) | [-0.57] |
| Log Likelihood | -887.2 | | | -282.4 | |
| Ps. R-squared | .017 | | | .001 | |
| LR $\chi^2$ | 30.79*** | | | 1.31 | |
| ♯ Firms | 6,367 | | | 790 | |
| ♯ Non Treated | 6,163 | | | 586 | |

Standard errors in parentheses. T-statistics in squared brackets produced from a Student t-test comparing the set of treated firms with the set of other firms. *** p<0.01, ** p<0.05, * p<0.1. $\bar{x}_{T=1}$: Mean of treated firms where $x \in \{\omega, y, k, l, m\}$; $\bar{x}_{T=0}$: Mean of non treated firms; $\bar{x}_{CF}$: Mean of firms included in the counterfactual.

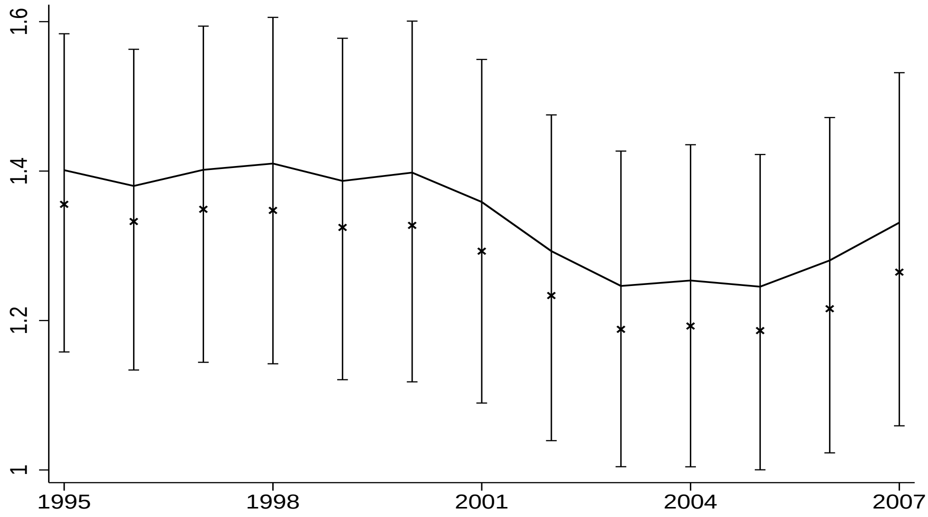


Figure 1: Evolution of markups between 1995 and 2007. The solid line, the marker "x" and vertical lines indicate the mean, median, and interquartile range, respectively.

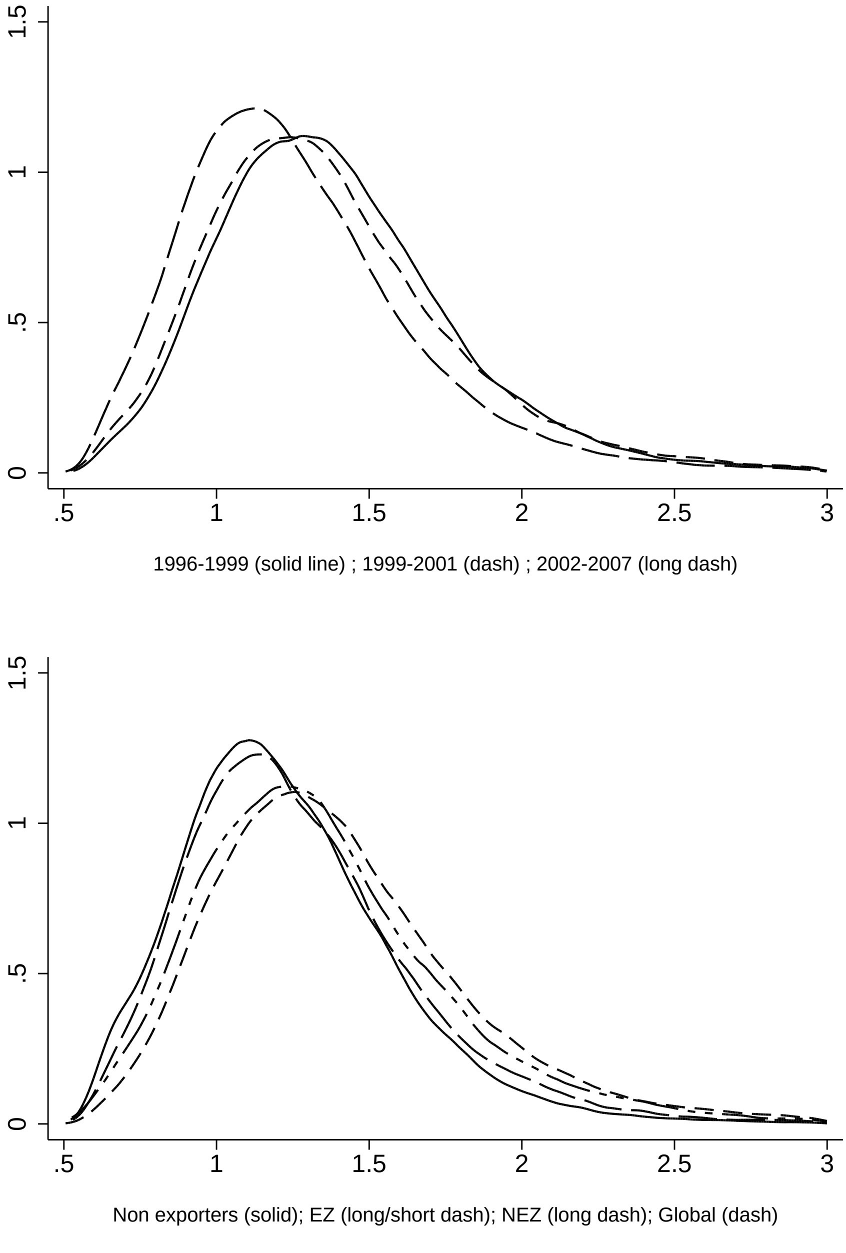


Figure 2: Distribution of Markups by Period (top panel) and by Destination of Customers (bottom panel)

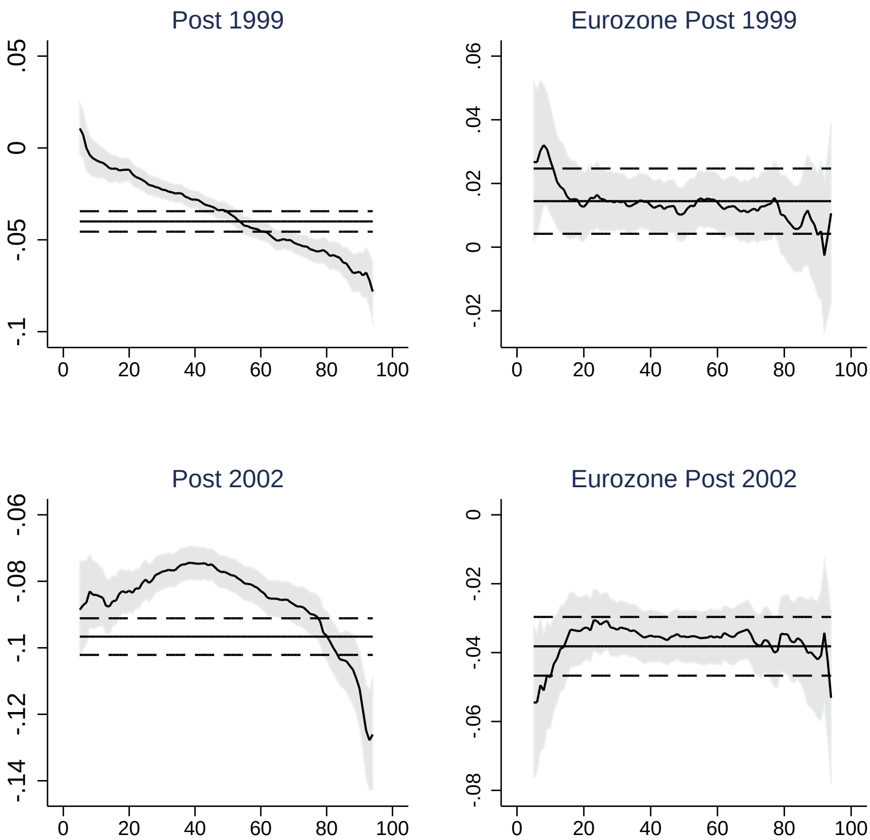


Figure 3: Marginal Effects of Export to the Eurozone per Quantile

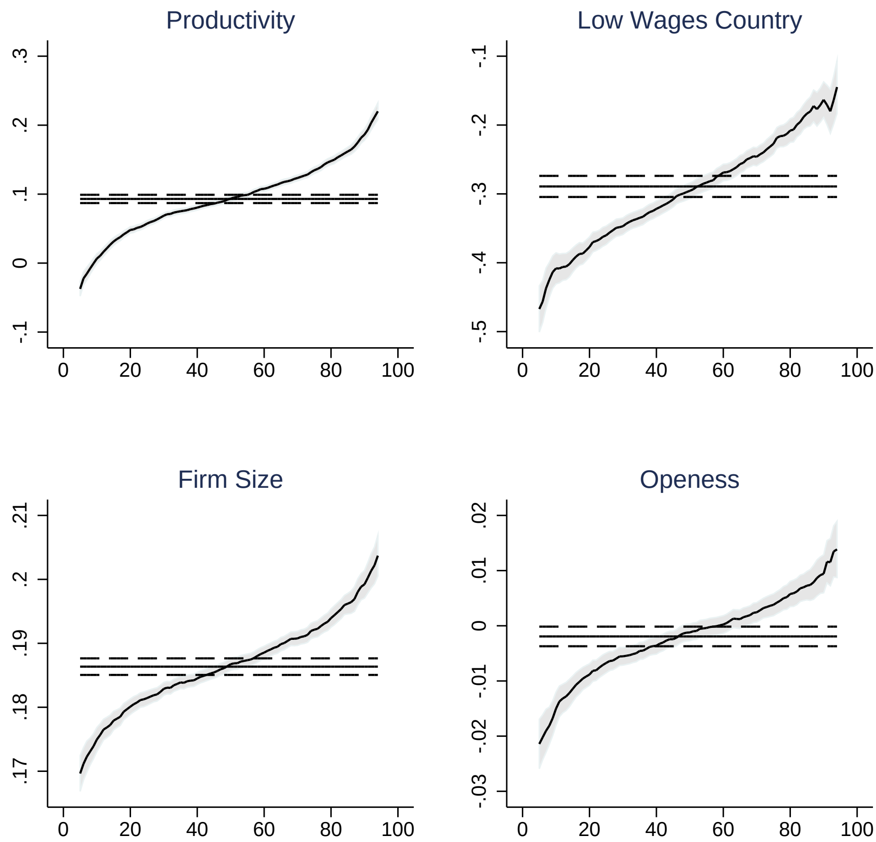


Figure 4: Marginal Effects of Firm and Industry Characteristics per Quantile

Table 1: Variable Used for the Estimation of Markup $\mu$ and Total Factor Productivity $\omega$, by Industry ($N = 215,049$)

| Industry | ♯ firms | $Y$ | $Q$ | $L$ | $K$ | $\hat{\mu}$ | $\hat{\omega}$ |
|---|---|---|---|---|---|---|---|
| All manufacturing | 29,193 | 8.923 | 7.944 | 11.58 | 7.323 | 1.339 | 1.478 |
| Automobile | 736 | 9.585 | 8.348 | 12.04 | 7.820 | 1.568 | 1.476 |
| Chemicals | 2,919 | 9.398 | 8.295 | 11.77 | 8.152 | 1.457 | 1.520 |
| Clothing & Footwear | 2,326 | 8.222 | 7.333 | 11.40 | 6.230 | 1.475 | 1.383 |
| Electric & Electronic Components | 1,223 | 9.190 | 8.263 | 11.86 | 7.462 | 1.383 | 1.420 |
| Electric & Electronic Equipment | 1,641 | 9.030 | 8.185 | 11.67 | 6.466 | 1.127 | 1.475 |
| House Equipment & Furnishings | 1,926 | 8.815 | 7.793 | 11.61 | 7.014 | 1.430 | 1.453 |
| Machinery & Mechanical Equipment | 5,065 | 8.820 | 7.858 | 11.48 | 6.928 | 1.213 | 1.500 |
| Manufacture of Coke and Fuel | 61 | 11.47 | 10.38 | 12.56 | 9.925 | 1.605 | 1.921 |
| Metallurgy, Iron and Steel | 4,993 | 8.658 | 7.767 | 11.43 | 7.450 | 1.283 | 1.450 |
| Mineral Industries | 1,201 | 9.030 | 8.003 | 11.63 | 7.950 | 1.408 | 1.468 |
| Pharmaceuticals | 756 | 10.26 | 9.094 | 12.27 | 8.247 | 1.435 | 1.571 |
| Printing & Publishing | 2,390 | 8.825 | 7.909 | 11.39 | 7.176 | 1.114 | 1.524 |
| Textile | 1,769 | 8.621 | 7.575 | 11.40 | 7.272 | 1.498 | 1.482 |
| Transportation Machinery | 426 | 9.502 | 8.568 | 12.16 | 7.716 | 1.364 | 1.373 |
| Wood & Paper | 1,752 | 9.017 | 7.953 | 11.56 | 7.784 | 1.488 | 1.508 |
| F-statistics | - | 1634.00 | 1490.96 | 758.50 | 2152.88 | 1840.16 | 475.05 |
| R-squared | - | .096 | .088 | .047 | .123 | .107 | .003 |

All F-statistics significant at 1% level. $Y$: Sales; $Q$: Value Added; $L$: Labor (hours worked); $K$: Capital Stocks. Variables $Y$, $Q$ and $K$ are expressed in 2000 constant euros. Variables $Y$, $VA$, $L$ and $K$ are log transformed.

Table 2: Participation Rates (PR), Export Intensity (EI), Entry Rates (ER) and Exit Rates (XR), by Export Market and Time Period

| Time Period | ♯ Obs. | | All | $Dest_{EZ}$ | $Dest_{NEZ}$ |
|---|---|---|---|---|---|
| All Periods 1995-2007 | 215,186 | PR | 72.8 | 57.7 | 67.9 |
| | | EI | 28.2 | 13.6 | 14.6 |
| | | ER | 5.0 | 4.9 | 6.3 |
| | | XR | 4.7 | 4.5 | 6.2 |
| Period 1995-1998 | 67,770 | PR | 71.6 | 54.8 | 69.0 |
| | | EI | 28.4 | 12.7 | 15.7 |
| | | ER | 4.7 | 5.0 | 5.2 |
| | | XR | 3.6 | 3.9 | 4.1 |
| Period 1999-2001 | 51,322 | PR | 73.9 | 60.3 | 67.1 |
| | | EI | 29.4 | 15.0 | 14.4 |
| | | ER | 5.3 | 6.2 | 6.7 |
| | | XR | 5.2 | 4.8 | 8.2 |
| Period 2002-2007 | 96,094 | PR | 73.2 | 58.4 | 67.6 |
| | | EI | 27.5 | 13.4 | 14.1 |
| | | ER | 5.1 | 4.1 | 6.9 |
| | | XR | 5.2 | 4.8 | 6.7 |

$D_{EZ}$: Eurozone markets; $D_{NEZ}$: Outside eurozone markets. The participation rate (PR) is defined as the number of firm-year exporting (to all markets, to the eurozone and to non-eurozone markets, respectively) over the overall number of firm-years. Export intensity (EI) displays exports values divided by sales, by destination markets. The entry rate (ER) is computed as the ratio of the number of firms new to exports (or new to the eurozone and to non-eurozone markets, respectively) over the number of firms exporting. The exit rate (XR) is computed as the ratio of the number of firms withdrawing from export markets (or from the eurozone and from non-eurozone markets, respectively) over the number of firms exporting.

Table 3: Computed Markups and Productivity Indices, by Time Period and Export Status

| Export Status | Time Period | Share | $\bar{\mu}$ | $\sigma_\mu$ | $\text{IQR}_\mu$ | $\bar{\omega}$ | $\sigma_\omega$ | $\text{IQR}_\omega$ |
|---|---|---|---|---|---|---|---|---|
| All firms | 1995-2007 | 100 | 1.339 | 0.388 | 0.456 | 1.478 | 0.270 | 0.272 |
| All firms | 1995-1998 | 100 | 1.398 | 0.379 | 0.442 | 1.421 | 0.254 | 0.256 |
| All firms | 1999-2001 | 100 | 1.381 | 0.404 | 0.467 | 1.455 | 0.267 | 0.261 |
| All firms | 2002-2007 | 100 | 1.274 | 0.376 | 0.439 | 1.530 | 0.273 | 0.276 |
| Non exporters (NX) | 1995-2007 | 27.1 | 1.251 | 0.337 | 0.401 | 1.460 | 0.223 | 0.231 |
| Non exporters | 1995-1998 | 28.4 | 1.310 | 0.325 | 0.386 | 1.409 | 0.207 | 0.226 |
| Non exporters | 1999-2001 | 26.1 | 1.281 | 0.352 | 0.403 | 1.427 | 0.226 | 0.220 |
| Non exporters | 2002-2007 | 26.8 | 1.192 | 0.329 | 0.381 | 1.515 | 0.220 | 0.234 |
| Eurozone only (EZ) | 1995-2007 | 5.0 | 1.345 | 0.389 | 0.470 | 1.513 | 0.260 | 0.251 |
| Eurozone only | 1995-1998 | 2.6 | 1.396 | 0.368 | 0.430 | 1.423 | 0.218 | 0.213 |
| Eurozone only | 1999-2001 | 6.8 | 1.380 | 0.399 | 0.490 | 1.477 | 0.242 | 0.235 |
| Eurozone only | 2002-2007 | 5.6 | 1.306 | 0.386 | 0.466 | 1.565 | 0.272 | 0.263 |
| Outside EZ (NEZ) | 1995-2007 | 15.1 | 1.289 | 0.363 | 0.433 | 1.488 | 0.232 | 0.237 |
| Outside EZ | 1995-1998 | 16.8 | 1.348 | 0.353 | 0.421 | 1.439 | 0.217 | 0.220 |
| Outside EZ | 1999-2001 | 13.6 | 1.321 | 0.377 | 0.442 | 1.470 | 0.225 | 0.224 |
| Outside EZ | 2002-2007 | 14.8 | 1.225 | 0.354 | 0.413 | 1.537 | 0.237 | 0.243 |
| Global (G) | 1995-2007 | 52.8 | 1.397 | 0.409 | 0.481 | 1.481 | 0.301 | 0.310 |
| Global | 1995-1998 | 52.1 | 1.463 | 0.403 | 0.472 | 1.423 | 0.288 | 0.293 |
| Global | 1999-2001 | 53.5 | 1.445 | 0.422 | 0.487 | 1.461 | 0.296 | 0.301 |
| Global | 2002-2007 | 52.8 | 1.326 | 0.395 | 0.460 | 1.532 | 0.304 | 0.316 |

IQR: Interquartile range

Table 4: Student $t$ and Kolmogorov-Smirnov Comparison Tests for Markups and Productivity Indices

| Comparison | $t$ | Conclusion | $D$ | Conclusion |
|---|---|---|---|---|
| | | Markup $\mu$ | | |
| EZ vs NX | 25.8 | $\bar{\mu}_{EZ} \geq \bar{\mu}_{NX}$ | $-0.102$ | $F(EZ) \geq F(NX)$ |
| EZ vs NEZ | 13.7 | $\bar{\mu}_{EZ} \geq \bar{\mu}_{NEZ}$ | $-0.061$ | $F(EZ) \geq F(NEZ)$ |
| EZ vs G | $-12.6$ | $\bar{\mu}_{EZ} \leq \bar{\mu}_{G}$ | 0.069 | $F(EZ) \leq F(G)$ |
| NEZ vs NX | 15.5 | $\bar{\mu}_{NEZ} \geq \bar{\mu}_{NX}$ | $-0.044$ | $F(NEZ) \geq F(NX)$ |
| NEZ vs G | $-43.3$ | $\bar{\mu}_{NEZ} \leq \bar{\mu}_{G}$ | 0.118 | $F(NEZ) \leq F(G)$ |
| G vs NX | 74.2 | $\bar{\mu}_{G} \geq \bar{\mu}_{NX}$ | $-0.160$ | $F(G) \geq F(NX)$ |
| P2 vs P1 | $-7.4$ | $\bar{\mu}_{P2} \geq \mu_{P1}$ | 0.043 | $F(P2) \leq F(P1)$ |
| P3 vs P1 | $-65.5$ | $\bar{\mu}_{P3} \geq \mu_{P1}$ | 0.126 | $F(P3) \leq F(P1)$ |
| P3 vs P2 | $-50.7$ | $\bar{\mu}_{P3} \geq \mu_{P2}$ | 0.128 | $F(P3) \leq F(P2)$ |
| | | Total Factor Productivity $\omega$ | | |
| EZ vs NX | 21.9 | $\bar{\omega}_{EZ} \geq \bar{\omega}_{NX}$ | $-0.070$ | $F(EZ) \geq F(NX)$ |
| EZ vs NEZ | 9.2 | $\bar{\omega}_{EZ} \geq \bar{\omega}_{NEZ}$ | $-0.038$ | $F(EZ) \geq F(NEZ)$ |
| EZ vs G | 10.5 | $\bar{\omega}_{EZ} \geq \bar{\omega}_{G}$ | $-0.096$ | $F(EZ) \geq F(G)$ |
| NEZ vs NX | 18.1 | $\bar{\omega}_{NEZ} \geq \bar{\omega}_{NX}$ | $-0.042$ | $F(NEZ) \geq F(NX)$ |
| NEZ vs G | 4.0 | $\bar{\omega}_{NEZ} \geq \bar{\omega}_{G}$ | $-0.087$ | $F(NEZ) \geq F(G)$ |
| G vs NX | 15.0 | $\bar{\omega}_{G} \geq \bar{\omega}_{NX}$ | $-0.075$ | $F(G) \geq F(NX)$ |
| P2 vs P1 | 21.7 | $\bar{\omega}_{P2} \geq \bar{\omega}_{P1}$ | $-0.060$ | $F(P2) \geq F(P1)$ |
| P3 vs P1 | 81.7 | $\bar{\omega}_{P3} \geq \bar{\omega}_{P1}$ | $-0.121$ | $F(P3) \geq F(P1)$ |
| P3 vs P2 | 51.0 | $\bar{\omega}_{P3} \geq \bar{\omega}_{P2}$ | $-0.144$ | $F(P3) \geq F(P2)$ |

All tests are significant at the 1% level: the Student t-test $H_0$: $\overline{x}_{G_1} = \overline{x}_{G_2}$, where $x$ stands for either markups $\mu$ or total factor productivity $\omega$, and the KS test for stochastic dominance of $F(G_1)$ over $F(G_2)$, where $F(\cdot)$ is the cumulative distribution of either markups $\mu$ or total factor productivity $\omega$. P1 stands for period 1995-1998, P2 stands for period 1999-2001, and P3 stands for period 2002-2007. NX, EZ, NEZ and G represent non-exporters, eurozone exporters, exporters outside the eurozone and global exporters, respectively.

Table 5: Definition of Variables and Data Sources

| Name | Variation | Definition | Source |
|---|---|---|---|
| $\mu$ | Firm-year | Firm Markups | EAE, Authors' Calculations |
| $\omega$ | Firm-year | Firm Productivity | EAE, Authors' Calculations |
| Sales | Firm-year | Firm Revenue (log) | EAE |
| LW Import | Sector-year | Low Wage Countries Import as a share of Total French Import per Industry | BACI-CEPII |
| Openness | Sector-year | Trade over Value Added per Industry | OECD STAN Database |
| Exporters | Firm-year | Dummy set to 1 when the French Firm is exporter | EAE |
| EZ only | Firm-year | Dummy set to 1 if firm $i$ exports to the Eurozone only | French Custom Data |
| NEZ only | Firm-year | Dummy set to 1 if firm $i$ exports outside the Eurozone only | French Custom Data |
| Global | Firm-year | Dummy set to 1 if firm $i$ exports both outside and inside the Eurozone | French Custom Data |
| p99 | Year | Dummy set to 1 if $t \geq 1999$ | |
| p02 | Year | Dummy set to 1 if $t \geq 2002$ | |

See previous table footnotes for acronyms.

Table 6: Descriptive Statistics

| Variable | Mean | Median | St.Dev. | Min. | Max. |
|---|---|---|---|---|---|
| Markup ($\mu$) | 1.339 | 1.280 | 0.388 | 0.206 | 4.337 |
| TFP ($\omega$) | 1.478 | 1.466 | 0.270 | -3.466 | 6.472 |
| Sales (logs) | 8.923 | 8.724 | 1.210 | 3.877 | 16.900 |
| LW Import Share | 0.086 | 0.052 | 0.090 | 0.002 | 0.387 |
| Openness | 2.728 | 2.253 | 2.228 | 0.570 | 43.230 |
| Exporters | 0.729 | 1.000 | 0.445 | 0.000 | 1.000 |
| EZ only | 0.050 | 0.000 | 0.217 | 0.000 | 1.000 |
| NEZ only | 0.151 | 0.000 | 0.358 | 0.000 | 1.000 |
| Global | 0.528 | 1.000 | 0.499 | 0.000 | 1.000 |

Table 7: Sequential OLS Regressions – Dependent Variable: Markup $\mu$ – 1995-2007

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Post 1999 | -0.030*** | -0.032*** | -0.027*** | -0.037*** |
| | (0.003) | (0.003) | (0.004) | (0.004) |
| Post 2002 | -0.124*** | -0.123*** | -0.123*** | -0.109*** |
| | (0.003) | (0.003) | (0.003) | (0.004) |
| TFP ($\omega$) | 0.143*** | 0.145*** | 0.145*** | 0.145*** |
| | (0.005) | (0.005) | (0.005) | (0.005) |
| EZ only | | 0.098*** | 0.084*** | 0.084*** |
| | | (0.004) | (0.009) | (0.009) |
| NEZ only | | 0.032*** | 0.034*** | 0.034*** |
| | | (0.002) | (0.004) | (0.004) |
| Global | | 0.144*** | 0.151*** | 0.151*** |
| | | (0.002) | (0.003) | (0.003) |
| EZ $\times$ Post 1999 | | | 0.016* | 0.008 |
| | | | (0.010) | (0.011) |
| NEZ $\times$ Post 1999 | | | -0.003 | -0.000 |
| | | | (0.005) | (0.007) |
| Global $\times$ Post 1999 | | | -0.010** | 0.008 |
| | | | (0.004) | (0.005) |
| EZ $\times$ Post 2002 | | | | 0.015 |
| | | | | (0.009) |
| NEZ $\times$ Post 2002 | | | | -0.004 |
| | | | | (0.006) |
| Global $\times$ Post 2002 | | | | -0.028*** |
| | | | | (0.005) |
| Exporters | 0.118*** | | | |
| | (0.002) | | | |
| Constant | 1.107*** | 1.106*** | 1.102*** | 1.102*** |
| | (0.007) | (0.007) | (0.007) | (0.007) |
| Observations | 215,049 | 215,049 | 215,049 | 215,049 |
| R-squared | 0.051 | 0.061 | 0.061 | 0.062 |

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. All regressions include a linear time trend.

Table 8: Firm-Level Markups and Export Status – 1995-2007

| | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|
| Post 1999 | -0.033*** | -0.036*** | -0.046*** | -0.068*** | -0.043** |
| | (0.003) | (0.004) | (0.006) | (0.005) | (0.020) |
| Post 2002 | -0.101*** | -0.096*** | -0.092*** | -0.104*** | -0.036* |
| | (0.003) | (0.003) | (0.006) | (0.007) | (0.018) |
| TFP ($\omega$) | 0.083*** | 0.114*** | 0.092*** | 0.187*** | 0.106 |
| | (0.014) | (0.017) | (0.026) | (0.034) | (0.073) |
| EZ × Post 1999 | 0.022*** | 0.031*** | 0.052* | 0.049* | 0.090* |
| | (0.008) | (0.009) | (0.028) | (0.027) | (0.050) |
| NEZ × Post 1999 | 0.002 | 0.001 | 0.014** | 0.016* | 0.034 |
| | (0.004) | (0.005) | (0.007) | (0.009) | (0.032) |
| Global × Post 1999 | 0.002 | 0.002 | 0.016*** | 0.019*** | 0.009* |
| | (0.004) | (0.004) | (0.006) | (0.006) | (0.019) |
| EZ × Post 2002 | 0.000 | -0.007 | -0.016 | -0.024 | -0.138** |
| | (0.006) | (0.007) | (0.031) | (0.027) | (0.058) |
| NEZ × Post 2002 | 0.004 | 0.007 | 0.004 | -0.000 | -0.056* |
| | (0.004) | (0.005) | (0.009) | (0.008) | (0.029) |
| Global × Post 2002 | -0.024*** | -0.025*** | -0.028*** | -0.031*** | -0.084*** |
| | (0.003) | (0.004) | (0.006) | (0.006) | (0.018) |
| Sales (logs) | 0.165*** | 0.176*** | 0.189*** | 0.187*** | 0.250*** |
| | (0.004) | (0.005) | (0.005) | (0.008) | (0.021) |
| LW imports | -0.291*** | -0.307*** | -0.287*** | -0.511*** | 0.029 |
| | (0.053) | (0.060) | (0.076) | (0.108) | (0.339) |
| Openness | -0.004 | -0.006 | -0.001 | -0.028** | 0.012 |
| | (0.005) | (0.006) | (0.010) | (0.014) | (0.024) |
| Constant | -0.134*** | -0.286*** | -0.390*** | -0.484*** | -0.932*** |
| | (0.038) | (0.047) | (0.061) | (0.086) | (0.211) |
| Observations | 215,049 | 148,428 | 82,137 | 20,655 | 8,166 |
| R-squared | 0.151 | 0.181 | 0.202 | 0.309 | 0.248 |
| Number of Firms | 29,178 | 12,515 | 6,908 | 6,908 | 790 |

Robust standard errors in parentheses in all regressions. Block bootstrap standard errors displayed in regression (8). *** $p<0.01$, ** $p<0.05$, * $p<0.1$. All regressions include a linear time trend. Regression (5) includes all observations. Regression (6) includes firms present at least from 1998 to 2003. Regression (7) include firms that do not change export status. Regression (8) collapses all firm-year variables into firm period (1995-1998; 1999-2001; 2002-2007). Regression (9) uses the sample of firms resulting from the matching procedure (see text).

Table 9: Firm-Level Markups and Intensity of Exports to the Eurozone – 1995-2007

| | (10) | (11) | (12) | (13) | (14) |
|---|---|---|---|---|---|
| Post 1999 | -0.034*** | -0.038*** | -0.040*** | -0.061*** | -0.058*** |
| | (0.003) | (0.003) | (0.004) | (0.004) | (0.015) |
| Post 2002 | -0.105*** | -0.099*** | -0.097*** | -0.110*** | -0.069*** |
| | (0.002) | (0.003) | (0.004) | (0.006) | (0.013) |
| TFP ($\omega$) | 0.083*** | 0.113*** | 0.093*** | 0.188*** | 0.105 |
| | (0.014) | (0.017) | (0.024) | (0.034) | (0.073) |
| $\alpha_{EZ}$ | 0.002 | 0.006 | 0.020** | 0.025** | 0.028 |
| | (0.005) | (0.005) | (0.010) | (0.013) | (0.027) |
| $\alpha_{EZ} \times$ Post 1999 | 0.006 | 0.010** | 0.014** | 0.015* | 0.039** |
| | (0.005) | (0.005) | (0.007) | (0.008) | (0.020) |
| $\alpha_{EZ} \times$ Post 2002 | -0.022*** | -0.027*** | -0.038*** | -0.038*** | -0.075*** |
| | (0.004) | (0.004) | (0.007) | (0.007) | (0.021) |
| Exports (logs) | 0.001** | 0.001*** | 0.003*** | 0.005*** | 0.004 |
| | (0.000) | (0.000) | (0.001) | (0.002) | (0.004) |
| Sales (logs) | 0.164*** | 0.174*** | 0.186*** | 0.182*** | 0.246*** |
| | (0.004) | (0.005) | (0.006) | (0.008) | (0.021) |
| LW imports | -0.299*** | -0.317*** | -0.289*** | -0.518*** | -0.122 |
| | (0.053) | (0.060) | (0.081) | (0.108) | (0.023) |
| Openness | -0.005 | -0.007 | -0.002 | -0.031** | -0.017 |
| | (0.005) | (0.006) | (0.008) | (0.014) | (0.023) |
| Constant | -0.130*** | -0.279*** | -0.407*** | -0.503*** | 0.968*** |
| | (0.038) | (0.046) | (0.063) | (0.085) | (0.214) |
| Observations | 215,049 | 148,428 | 82,137 | 20,655 | 8,166 |
| R-squared | 0.150 | 0.180 | 0.202 | 0.309 | 0.248 |
| Number of Firms | 29,178 | 12,515 | 6,908 | 6,908 | 790 |

Robust standard errors in parentheses in all regressions. Block bootstrap standard errors displayed in regression (12). *** p<0.01, ** p<0.05, * p<0.1. All regressions include a linear time trend. Regression (10) includes all observations. Regression (11) includes firms present at least from 1998 to 2003. Regression (12) includes firms that do not change export status. Regression (13) collapses all firm-year variables into firm period (1995-1998; 1999-2001; 2002-2007). Regression (14) uses the sample of firms resulting from the matching procedure (see text).

Table 10: Fixed Effect Quantile Regressions – Dependent Variable: Firm-Level Markup $\breve{\mu}$ – Restricted Sample 1995-2007, $N = 82,137$

| | (12) LSDV | (15) $\tau = 0.1$ | (16) $\tau = 0.25$ | (17) $\tau = 0.5$ | (18) $\tau = 0.75$ | (19) $\tau = 0.9$ |
|---|---|---|---|---|---|---|
| Post 1999 | -0.040*** | -0.007 | -0.019*** | -0.035*** | -0.055*** | -0.067*** |
| | (0.003) | (0.005) | (0.003) | (0.002) | (0.002) | (0.006) |
| Post 2002 | -0.097*** | -0.084*** | -0.079*** | -0.078*** | -0.090*** | -0.112*** |
| | (0.003) | (0.005) | (0.002) | (0.003) | (0.003) | (0.005) |
| TFP ($\omega$) | 0.093*** | 0.007* | 0.057*** | 0.094*** | 0.135*** | 0.186*** |
| | (0.003) | (0.004) | (0.002) | (0.003) | (0.003) | (0.004) |
| $\alpha_{EZ}$ | 0.020*** | 0.006 | 0.012*** | 0.020*** | 0.027*** | 0.037*** |
| | (0.004) | (0.006) | (0.004) | (0.003) | (0.004) | (0.010) |
| $\alpha_{EZ}\times$ Post 1999 | 0.014*** | 0.027*** | 0.015*** | 0.011*** | 0.013*** | 0.004 |
| | (0.005) | (0.007) | (0.005) | (0.003) | (0.005) | (0.013) |
| $\alpha_{EZ}\times$ Post 2002 | -0.038*** | -0.047*** | -0.032*** | -0.035*** | -0.037*** | -0.042*** |
| | (0.004) | (0.006) | (0.003) | (0.004) | (0.004) | (0.008) |
| Exports (logs) | 0.003*** | 0.004*** | 0.004*** | 0.003*** | 0.002*** | 0.002*** |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Sales (logs) | 0.186*** | 0.175*** | 0.181*** | 0.187*** | 0.192*** | 0.199*** |
| | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| LW Import | -0.289*** | -0.408*** | -0.360*** | -0.296*** | -0.227*** | -0.163*** |
| | (0.008) | (0.013) | (0.008) | (0.006) | (0.010) | (0.013) |
| Openness | -0.002** | -0.015*** | -0.007*** | -0.001 | 0.004*** | 0.009*** |
| | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| Constant | -0.407*** | -0.348*** | -0.390*** | -0.413*** | -0.438*** | -0.479*** |
| | (0.008) | (0.007) | (0.007) | (0.008) | (0.007) | (0.011) |

Bootstrap standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. All regressions include a linear time trend. The restricted sample includes firms present at least from 1998 to 2003 that do not change export status. Column (15) displays the results of a fixed effect least squares regression (R-squared = 0.713).