The Exchange Rate, Employment and Hours: What Firm-Level Data Say*

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Abstract

Using a representative panel of manufacturing firms, we estimate the response of job and hours worked to currency swings, showing that it depends primarily on firms' exposure to foreign sales and their reliance on imported inputs. We also show that, for given international exposure, the response to exchange rate fluctuations is magnified when firms exhibit a lower monopoly power and when they face foreign pressure in the domestic market through import penetration. The degree of substitutability between imported and other inputs and the distribution of workers by type introduce additional degrees of specificity in the employment sensitivity to exchange rate swings. We show that firms' export status and episodes of entry and exit in the export market are associated with a heterogeneous employment response to exchange rate variations. Wage adjustments are shown to provide a channel through which firms react to currency shocks. Finally, gross job flows within the firm are found to react to exchange rate fluctuations, but the effect on job creation is predominant.

JEL classifications: E24; F16; F31.

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1 Introduction

Do exchange rate fluctuations affect employment and decisions about hours worked across firms? While there is a substantial literature on the implications for the real economy of exchange rate swings, and at least some contributions focusing on the effects on labor market variables, there is little evidence at the microeconomic level on the employment and hours response to currency movements. In this paper we investigate this relationship in detail, analyzing the transmission channels at the firm level of an exchange rate shock to jobs and hours worked.

There are at least two major reasons why focusing on firm level data is particularly appropriate for studying the relationship between employment and exchange rates. First, this level of disaggregation permits a better understanding of the transmission mechanisms at the heart of this relationship, compared to what emerges from more aggregate data. Second, it allows us to appraise how exchange rate fluctuations can have heterogenous effects on firms with different characteristics.

The literature in this area can be broadly divided into two distinct groups. The first set of contributions deals with the influence of exchange rate variations and, in general, international factor prices on net employment, either at aggregate or at industry level (Branson and Love, 1988; Revenga, 1992; Burgess and Knetter, 1998; Goldberg and Tracy, 2000; Campa and Goldberg, 2001). A more recent body of theoretical and empirical research, based on the flow approach to labor markets, investigates instead the impact of the exchange rate on gross job flows and, more generally, on the process of inter- and intra-industry employment reallocation (Gourinchas, 1998 and 1999; Klein et al., 2002 and 2003).

In our study we investigate a variety of channels through which the currency value affects firms' labor demand and net employment. Following Campa and Goldberg (2001), we characterize the direction and magnitude of labor response as primarily depending on the producers' external orientation through both exports and imported inputs use. However, unlike their study, we find a statistically significant and non-negligible effect of exchange rate movements on both employment and hours worked. This effect is seen separately on the revenue side, through exposure to foreign sales, and on the cost side, through reliance on imported inputs. It is our belief, one which we support with consistent empirical evidence, that Campa and Goldberg's finding of a low degree of responsiveness of labor input to exchange rate fluctuations, and thus the divergence from our own results, depends on their use of data aggregated at the industry level. Indeed, a large volume of

gross job flows has been widely documented also within narrowly defined industries. As a result, an appraisal of the effect of exchange rates on *aggregate* net employment would easily hide the effects on intra-sector job reallocation (Gourinchas, 1999). More generally, recent work on labor decisions underline the importance of firms' individual characteristics, which are difficult to track in the aggregate data (see, for example, Davis and Haltiwanger, 1999). This is why we analyze the responsiveness of employment and hours worked to exchange rate fluctuations using firm-level panel data obtained from two high-quality sources: the Bank of Italy Survey of Investment in Manufacturing and the Company Accounts Data Service reports. The period we analyze is antecedent to the introduction of the single European currency and provides a very interesting case study, as Italy experienced significant exchange rate oscillations for a developed country, and at the same time was characterized by a high degree of firms' external orientation on both the revenue and cost sides.

From a theoretical point of view, the different channels through which exchange rate variations affect firms' decisions on labor input can be easily described: after an exchange rate swing – let's say a depreciation – the more a firm relies on imported inputs, the larger the increase in its costs, the reduction in marginal profitability, and the ensuing drop in employment and hours worked. At the same time, in the aftermath of a depreciation the higher the firm's external sale exposure – i.e., the share of export revenues in total revenues – the larger the increase in its sales, in marginal profitability and therefore in the use of labor input.

For an imperfectly competitive firm, anther important theoretical implication is that the sensitivity of employment and hours worked to exchange rate fluctuations depends on a firm's degree of monopoly power. There are two channels through which market power affects the relationship. The first is related to the extent to which firms "pass-through" an exchange rate shock into export prices expressed in foreign currency. In particular, all other things being equal, the lower the exchange rate pass-through into destination market prices – i.e., the more stable local-currency prices are – the higher the labor inputs sensitivity to currency shocks. Indeed, as shown in a number of contributions analyzing the relationship between exchange rates and prices (see, e.g., Dornbusch, 1987; Knetter, 1993; Yang, 1997), the market structure affects the degree of exchange rate pass-through into export prices denominated in foreign currency. In particular, the pass-through is shown to be more pronounced if product differentiation is high, if the degree of substitution among different variants of goods is low and, in general, if the exporter faces a low degree of competition in the foreign market, thus having a higher pricing power. The second channel is related to the price elasticity of demand. For a given (non-zero) exchange-rate pass-through to destination market prices, a depreciation leads to a reduction of the export price in foreign currency, inducing a rise

in volumes sold in the foreign market (and, therefore, in profitability and labor demand), which is positively related to the price elasticity of foreign demand faced by firms. Since this elasticity is inversely related to a firm's degree of market power, the latter introduces an additional source of heterogeneity in the labor response to the exchange rate. In other words, for a given firm's external orientation, the effect of exchange rate swings on employment is magnified if firms exhibit low market power.

Our empirical findings are consistent with the theoretical predictions that the sensitivity of employment and hours worked to exchange rate fluctuations is higher for firms with low market power. Moreover, in the empirical analysis we single out three additional firm-specific features that are potentially relevant for characterizing the exchange rate-labor input link at the firm level. First, we consider the degree of import penetration in the domestic market where the firm operates. In industries where imports account for a large share of total demand, firms with a higher dependence on domestic revenues are more exposed to foreign competition and therefore to the effects of exchange rate swings. In other words, if import penetration is high in a given industry, then a currency appreciation would severely reduce the competitiveness of domestic firms, even more so for firms with a high internal orientation of their sales. Second, we consider the degree of substitutability in the production function between imported and domestically produced inputs. If technological features or market constraints prevent imported inputs from being substituted with other inputs, then an exchange rate swing is likely to have a more pronounced impact on employment and hours worked, as firms profitability will be more deeply affected by the external shock. Finally, we study whether the job and hours response to currency movements may depend on the distribution of workers by type of job within each firm (i.e., blue- vs. white-collar). This would clearly be the case if the firm's ability to adjust its labor force hinged on the type of workers it employs.

Perhaps more importantly, in light of the substantial microeconomic heterogeneity among firms in the decision to export (see, for example, Bernard et al., 2007), we investigate whether differences in firms' export status lead to a different employment response to an exchange rate shock. In so doing, we conduct a separate analysis of firms that are currently exporting, non-exporters, and firms that switch export status, from non-exporter to exporter and from exporter to non-exporter. As expected, we find that exchange rate swings affect employment differently in each of these four cases.

Our analysis also focuses on the wage response to exchange rate movements. The theoretical model provides some testable implications in this regard. Whilst our results for the response of employment and hours worked are quite different from those of Campa and Goldberg (2001), when we turn to investigating the wage response, the findings are more similar, with a statistically

significant influence of exchange rate movements on wages. As we discuss in more detail below, the fact that a significant wage response is also found at the industry level does not diminish the importance of job reallocation within industries. Indeed, following a currency movement, some jobs can be lost and others can be created within the same industry. However, if the wage packages offered to those taking the new jobs are different from those characterizing the jobs that have been lost, then there are significant wage changes at the industry level even without controlling for labor reallocation.[1]

A further characterization of the transmission of exchange rate shocks to firm level employment deals with gross job flows. Unlike the literature mentioned above on the impact of exchange rate on inter- and intra-industry job reallocation (Gourinchas, 1998 and 1999; Klein et al., 2002 and 2003), in our analysis we simply verify whether the net employment change following an exchange rate swing is accompanied by a considerable degree of job reallocation *within* the firm. Our results show that exchange rate swings have a significant effect on the rate of job reallocation within each firm. However, we also provide evidence that this effect is driven primarily by the exchange rate impact on job creation rather than on job destruction.

The remainder of the paper is organized as follows: section 2 presents the theoretical background of our analysis and the empirical specification; section 3 describes the data used; section 4 documents the empirical results from the baseline specification, and section 5 provides some further characterizations of the link between exchange rate and labor variables, dealing, for example, with firms' trading status, the response of wages and the implications for job reallocation within each firm. Section 6 concludes. The appendix presents a simple model providing a theoretical background for our empirical analysis.

2 Theoretical background and empirical specification

The mechanisms through which exchange rate variations affect employment levels can be ascertained in a simple framework, where labor demand is obtained from the first order conditions for a firm's profit maximization in an imperfectly competitive market and labor supply is assumed to be a simple function of wages and aggregate demand conditions (Campa and Goldberg, 2001). Exchange rate variations affect labor demand through the changes induced in the firm's marginal revenue product of labor via both the cost and the revenue channels. The effect on the cost side depends on a number of firm-specific characteristics, most notably its reliance upon imported in-

[1] Along these lines, Campa and Goldberg (2001) argue convincingly that the problems of aggregation are less apparent in estimating real wage elasticities than they are in estimating employment elasticities.

puts and the elasticity of substitution between these inputs and domestically produced substitutes. The effect on the revenue side depends primarily on the share of export revenues with respect to total revenues, the firm's market power in the product market, and the degree of exchange rate pass-through into export prices expressed in the foreign currency.

As is shown in the appendix, the elasticity of equilibrium employment, $\widetilde{L}$, with respect to the exchange rate is given by:

$$\frac{\Delta \widetilde{L}}{\Delta e}\frac{e}{\widetilde{L}} = \frac{1}{\overline{\mu}\beta}[(1-\chi)\eta_{p,e} - \chi(1-\eta_{p^*,e}) + \alpha(1-\eta_{s^*,e})]\frac{a_1}{1+a_1}, \tag{1}$$

where e is the level of the exchange rate, expressed as the number of foreign currency units per domestic currency unit (so that a decrease in the exchange rate amounts to a depreciation); $\chi \in [0,1]$ is the share of sales on foreign markets in total sales; $\alpha \in [0,1]$ is the share of production costs on imported inputs in total costs; $\overline{\mu}$ is the average value of the mark-up in the home and foreign markets; β is the share of labor costs over total revenues; $\eta_{p,e} \in [-1,0]$ and $\eta_{p^*,e} \in [0,1]$ are the elasticities of, respectively, domestic and foreign prices with respect to the exchange rate (i.e., the pass-through elasticities); $\eta_{s^*,e} \in [0,1]$ is the elasticity of foreign input prices with respect to the exchange rate; a_1 is a parameter describing the wage elasticity of labor supply.

2.1 The propagation mechanism

Equation (1) provides a number of testable insights into the mechanisms underlying the employment response to exchange rate variations. A first implication is that a firm's external orientation towards international markets is pivotal in shaping the direction and size of the effect on employment and hours worked of exchange rate swings. The transmission of exchange rate movements is seen through its effects on the marginal revenue product of labor and, therefore, on profits. It is important to emphasize that the expression $(1-\chi)\eta_{p,e} - \chi(1-\eta_{p^*,e})$ in the previous equations is clearly non-positive, given that $\chi \in [0,1]$, $\eta_{p,e} \in [-1,0]$ and $\eta_{p^*,e} \in [0,1]$. Therefore, an exchange rate depreciation (i.e., a decrease in e) has a positive effect on employment through the revenue channel. Moreover, the higher the share of foreign sales in total sales, χ, the stronger the increase in employment induced by an exchange rate depreciation.[2] By contrast, the expression $\alpha(1-\eta_{s^*,e})$ is non-negative, since $\eta_{s^*,e} \in [0,1]$, indicating that the same depreciation has a negative effect on employment through the cost channel. By the same token, the higher a firm's reliance on imported inputs relative to total input purchases, α, the more sizeable the employment decrease following an exchange rate depreciation.

[2] A necessary and sufficient condition for the magnification effect of foreign exposure to hold is that the sum of the (absolute values of) exchange rate pass-through elasticities be less than one (i.e., $|\eta_{p,e}| + \eta_{p^*,e} < 1$).

In addition to these basic elements characterizing a firm's orientation to foreign markets, two important supplementary features affect the sign and size of the exchange rate elasticity of employment and hours worked. The first deals with the degree of exchange rate pass-through that may influence the employment sensitivity to currency swings. The pass-through elasticity in the foreign market, $\eta_{p^*,e}$ – ranging from zero (no pass-through) to one (complete pass-through) – measures how willing the firm is to change the prices set in the foreign market as a result of exchange rate variations. It is clear from equation (1) that the lower the pass-through of exchange rate into foreign-currency export prices, $\eta_{p^*,e}$, the stronger the employment and hour response to exchange rate fluctuations. The relationship between exchange rates and prices have been widely analyzed and many studies have focused primarily on how the degree of exchange rate pass-through to export prices is determined by the market structure. A notable finding in the theoretical and empirical literature is that the degree of pass-through is less pronounced – i.e., foreign-currency prices tend to be more stable – when products are scarcely differentiated and the extent of their substitution among different variants is high (Yang, 1997). In general, the pass-through tends to be low if the exporting firm faces a high degree of competition in the foreign market and consequently has a limited pricing power (see, for example, Dornbusch, 1987; Knetter, 1993). Clearly, if the foreign market is perfectly competitive, so that the firm is a price taker, the pass-through elasticity is null and an exchange rate appreciation has a one-to-one effect on the value of foreign sales expressed in the domestic currency.

The second insight from equation (1) concerns the higher sensitivity of employment to exchange rate variations in firms exhibiting a lower market power, as measured by the mark-up index, $\overline{\mu}$. This results suggests that, for a given degree of foreign exposure, χ, and for a given (non-zero) value of the exchange rate pass-through elasticity, $\eta_{p^*,e}$, the reduction in the foreign currency export price induced by an exchange rate depreciation generates an increase in foreign sales, q^*, – and, as a consequence, of profitability and labor input – which is larger the higher is the price elasticity of foreign demand. Since, as it is well known, such elasticity is negatively related to a firm's degree of market power, the sensitivity of employment to exchange rate variations is amplified when a firm's market power is low.

In addition, by making foreign products less expensive, a currency appreciation decreases the competitiveness of domestic firms, causing a reduction of its sales in the home market, and therefore a drop in its profitability and employment. Again, market structure plays a crucial role in this respect, and the mechanism is similar to the one described above. The domestic pass-through elasticity ranges from minus one (complete pass-through) to zero (no pass-through), and (in absolute value) it is a decreasing function of the firm's market power in the domestic market. In the extreme

case of perfect competition in the domestic market, the competitive pressure from foreign producers determined by a currency appreciation must be rebated by a one-to-one price reduction by the domestic firm (i.e., the pass-through elasticity is minus one). Hence, a currency appreciation causes a reduction in the value of domestic sales, and the lower the firm's monopoly power the steeper the decline. Moreover, as explained convincingly by Dornsbusch (1987) and Campa and Goldberg (2001), the domestic pass-through elasticity is proportional to the degree of import penetration, which summarizes the amount of pressure in the domestic market exerted by foreign producers. This dependence will also be considered in our empirical analysis.

Finally, similar to the revenue side, the effect of an exchange rate depreciation on employment coming from an increase in the cost of imported inputs also depends on the degree of competition in the market for inputs. The closer to zero the elasticity of foreign input prices to the exchange rate, $\eta_{s^*,e}$ – which ranges from zero (no pass-through) to one (complete pass-through) – the larger the effect of exchange rate variations on employment.

2.2 Empirical specification and estimation

Based on the testable implications provided by our theoretical framework, for the econometric analysis we specify the following dynamic equations for labor input:

$$\Delta n_{it} = \beta_0 + \beta_1 \alpha_{it-1} \Delta e_t + \beta_2 \chi_{it-1} \Delta e_t + \beta_3 \alpha_{it-1} + \beta_4 \chi_{it-1} + \\ + \beta_5 \Delta s_{it-1} + \beta_6 MKUP_{it-1} + \beta_7 \Delta n_{it-1} + b' Z_{it} + \lambda_i + u_{it}, \quad (2)$$

which is the empirical counterpart of equation (1) and equation (A11) in the appendix. Given the non-stationarity of the exchange rate variable, we choose a specification in first-differences. Lower-case letters refer to the logarithmic transformation of the variable; N_{it} is the level of labor input of firm i at time t, S_{it} denotes the value of real sales and E_t is the real effective exchange rate, expressed as before in terms of number of foreign currency units per unit of domestic currency, so that an increase in the exchange rate amounts to an appreciation. $MKUP_{it}$ is a measure of firms' market power and Z_{it} is a vector of dummy variables referring to the year, the industry, the size and the geographic location of each firm. As labor input measures, we alternatively use the number of employees and the number of hours worked.

To appraise the implications for firms' labor variables of exchange rate swings, the key elements are the two interaction terms of the exchange rate change with, respectively, α_{it-1}, the share of costs for imported inputs in total variable costs, and χ_{it-1}, the share of export sales in total sales, both lagged by one period to mitigate the possible simultaneity bias arising from the effect of exchange

rate variations on the degree of firms' external orientation. As illustrated in the previous sections, these two variables reflect firms' exposure to international competition in foreign input markets (cost side) and foreign product markets (revenue side), respectively. By interacting the exchange rate changes, Δe_t, with the firm-specific, time-varying external orientation variables (α_{it-1} and χ_{it-1}), our empirical specification allows the estimated effect of the exchange rate on labor input to vary over time and across firms, depending on the two variables shaping the extent of firms' foreign exposure. In addition to the interaction terms, the imported input and export shares are also inserted in the specification as single regressors. Without the inclusion of these two variables in isolation, it would be difficult to ascertain whether a significant coefficient on the interaction terms reflects the effect of firms' external orientation rather than the effect of exchange rate movements.

Following Campa and Goldberg (2001), we single out more precisely the relevant effects of the currency value in the interaction terms, $\alpha_{it-1}\Delta e_t$ and $\chi_{it-1}\Delta e_t$, using the permanent component of the (log) change of, respectively, the import and export exchange rates, E_t. The reason for using the permanent components is that the theoretical model presented in the appendix can be seen as nested in a more general multi-period, stochastic firm's optimization problem. Under the two hypotheses that the unique source of uncertainty in the model is the exchange rate and that the latter follows a random walk, it has been shown that all model predictions from a multi-period set-up, which are relevant for our research purposes, can be appropriately obtained and investigated in a simpler framework such as the one devised in this paper (Campa and Goldberg, 1999 and 2001; Nucci and Pozzolo, 2001). If the exchange rate follows a random walk, then the conditional expected value of its future level is equal to today's realization and the effects of any exchange rate variation are persistent. Therefore, for establishing a closer link between this implicit assumption of the model and the empirical analysis, it is appropriate to focus on the permanent component of exchange rate fluctuations. On the other hand, as described in more detail in section 3, using a Beveridge and Nelson decomposition to identify the permanent component, we find that the residual transitory component of changes in the currency value accounts for only a small fraction of the variance of actual exchange rate variations. In fact, our empirical results are substantially unchanged if the actual values of exchange rate variations, instead of their permanent component, are used.

A lagged value of the dependent variable is included in the equation to control for the adjustment process typical of employment changes. The dynamic adjustment in our specification in first-difference is consistent with the predictions of the theoretical model, when the latter is augmented with the introduction of adjustment costs (see the appendix). We also include the change in the value of firms' real sales to account for demand conditions, and the lagged value of the mark-

up, to account for the effect of marginal profitability that is independent of the exchange rate shock. In addition to year dummies, controlling for time-varying effects common to all firms, the empirical specification allows for a firm-specific latent heterogeneity by including fixed effects, λ_i. The error terms of the specification above, u_{it}, are assumed to have finite moments with $E(u_{it}) = E(u_{it}u_{is}) = 0$, for all $t \neq s$.

In estimating the baseline equation (2), as well as a number of extensions, we use the generalized method of moments (GMM) estimator developed for dynamic panel data model (Arellano and Bond, 1991; Arellano and Bover, 1995; Blundell and Bond, 2000). The reason we rely on this methodology is that, in the specification above, the lagged values of the dependent variable, of real sales, and of the mark-up are correlated with the individual fixed effects, λ_i. This would yield a specification error, inducing inconsistency of the parameters obtained with standard panel estimators. The GMM technique controls for the endogeneity of the regressors and restores consistency of parameters' estimates. Moreover, this estimator turns out to be efficient within the class of instrumental variable estimators. In particular, following Arellano and Bover (1995) and Blundell and Bond (2000), we use the system GMM panel estimator.[3] Consistent with this approach, as GMM-type of instruments we selected the lagged values of the dependent variable, of real sales, and of the mark-up dated period t-2 and earlier. We also test for instruments exogeneity and, in general, for the robustness of the specification by performing the Hansen test of over-identifying restrictions, which allows us to verify the orthogonality between instruments and errors. Finally, we perform the Arellano-Bond test for second-order serial correlation based on residuals from the transformed equation (Arellano and Bond, 1991).

In the next section, we present the data set used in the empirical analysis and in section 4 we document and discuss the results.

3 Data and summary statistics

Our empirical analysis is conducted using firm level data obtained from the Survey of Investment in Italian Manufacturing (SIM) and from the Company Accounts Data Service reports. The Survey of Investment has been conducted every year since 1984 by the Bank of Italy on a stratified sample

[3]This estimator augments the Arellano and Bond (1991) estimator by building a system of two equations: the original equation and a transformed one. As in Arellano and Bond (1991), in the transformed equation a variety of instruments in levels can be used. However, under the novel approach a further assumption is made: that first differencing the instrumenting variables in the original equation make them uncorrelated with fixed effects. This allows us to exploit an even larger number of orthogonality conditions than before, by resorting to a larger instrument set.

of over 1,000 firms with more than 50 employees. It collects information at the firm level on a wide set of economic variables, including total employment, hours worked, total revenues and revenues from foreign sales. The Company Accounts Data Service reports contain detailed balance-sheet information on a sample of around 40,000 firms. Merging the information from the two sources resulted in an unbalanced panel of slightly less than 2,400 firms.[4] The empirical specifications are estimated using these microeconomic data on the period 1984 - 1998, i.e. the one preceding the introduction of the euro.

Because neither of our data sources provide distinct information on domestically produced and imported non-labor inputs, as in Nucci and Pozzolo (2001) we supplement the available firm-level information on purchases of materials and services and on labor costs with data from the 44-sectors input-output table of 1992 for the Italian economy. In particular, for each industry we select the values of intermediate inputs imported and also of all inputs purchased, both domestically produced and imported. We then use economic time series on import demand and production for each industry to update (backward and forward) the corresponding figures of the input-output table. Finally, in computing α_{it}, the share of expenditure on imported inputs in total input purchases of firm i at time t, we combine the industry-level information with data on firms' expenditure for materials and services and for labor inputs.[5] Admittedly, our measure of imported input share falls short of accurately reflecting the within-industry heterogeneity of firms' import behavior, so that the ensuing problem of imperfect measurement is likely to introduce an attenuation bias in the coefficient estimate.

Each firm's market power is computed as the ratio of profit margins to unit price. As suggested by Domowitz et al. (1986), we adopt the following time-varying measure:[6]

$$MKUP_{it} = \left(\frac{\text{Sales} + \text{Change in Inventories} - \text{Payroll} - \text{Cost of non-Labor Inputs}}{\text{Sales} + \text{Change in Inventories}} \right)_{it} \qquad (3)$$

[4] In order to verify that the survival rate of firms in the sample is not affected by exchange rate movements, we estimated a logit model on our data, where the firm-level dependent variable takes the value of one if the firm exited from the sample in the previous year and zero if the firm is in the sample in a given year. The explanatory variables are the one-period lagged exchange rate variation and a set of control dummy variables. The regression results indicate that the estimated coefficient of the exchange rate is not statistically different from zero.

[5] In particular, the firm-level share of expenditure on imported inputs in total expenditure is defined as $\alpha_{it} = \frac{\left(\frac{IM_{jt}}{TE_{jt}}\right) TE_{it}}{TE_{it}+LC_{it}}$, where IM_{jt} is the value of imported non-labor inputs by industry j (the sector to which firm i belongs), TE_{it} and TE_{jt} are the values of non-labor expenditure for intermediate inputs of, respectively, firm i and industry j, and LC_{it} is labor costs of firm i.

[6] The empirical measure of mark-up defined above has a strict relationship with the index discussed in the theoretical section: $MKUP = \frac{(\overline{\mu}-1)}{\overline{\mu}}$, assuming no changes in inventories.

Due to a lack of information on foreign profits, we cannot distinguish between the mark-ups in the home and foreign markets. A notable feature of the Bank of Italy's Survey is that, in addition to the number of employees and hours worked, it provides information on gross job flows at the firm level and, in particular, on the number of jobs created and of jobs lost in every year. As we show in the following section, this enables us to compute the rates of job creation and job destruction and of job reallocation within the firm and investigate their additional information content with respect to standard measures of net changes in employment and hours.

We measure currency oscillations using the export and the import real effective exchange rates of the lira computed by the Bank of Italy, each one taking into account 24 bilateral exchange rates. For the export rate, the weights reflect: a) the shares of Italian exports towards each of the 24 countries; b) the weight of each country's exports in satisfying the domestic demand of all 24 trade partners; c) the domestic demand that originates from each country. For the import rate, the weights simply reflect the geographic composition of Italian imports. The two real exchange rates are computed using producer price indexes (Banca d'Italia, 1998). In presence of even small adjustment costs, firms are unlikely to modify their employment level after temporary fluctuations in the exchange rates. For this reason, we derived the permanent component of exchange rate variations using the Beveridge and Nelson (1981) procedure that decomposes an I(1) time series into its permanent and transitory components. Since the real exchange rate we use are trade-weighted multilateral indexes, following Campa and Golberg (2001) we relied on a univariate decomposition procedure. In particular, we selected an autoregressive model for both the export and import monthly (log) real exchange rate change. From our decomposition, the transitory component accounts for only a small fraction of the variance of the actual change in the exchange rate. Consistent with the specification in the theoretical model, a reduction in the exchange rate is a (real) depreciation. Figure 1 shows the time profile of the monthly data on import and export real exchange rates in the period analyzed. While they exhibit a very similar pattern, some differences emerge in the mid-eighties and at the beginning of the nineties.

Table 1 documents some descriptive statistics for the variables used in our empirical analysis. The average rate of annual changes of the import exchange rate (its permanent component) is -0.29 per cent during the sample period, but the standard deviation is 5.40 per cent and the values at the first and third quartiles are -2.29 and 3.07 per cent, respectively, suggesting a non-negligible variability. In the case of the export exchange rate, the mean rate of change is -0.09 per cent, but the value of the standard deviation points to an even larger variability. The average degree of a firm's dependence on imported inputs is 13.9 per cent, with a distribution slightly skewed to the right, as suggested by the lower median level. The average share of export revenues in total

revenues, 29.8 per cent, is more than double that of import dependence, and also in this case its distribution is skewed to the right. Interestingly, even at the 25th percentile of the distribution, the share of export revenues is non-negligible (4.9 per cent). As we analyze in the section below on the firm extensive margin of export, arguably this reflects to a considerable extent the fact that our dataset deals with firms with more than 50 employees. More generally, however, while this finding is at odds with the evidence reported for the U.S. and the U.K. that exporting activity is relatively uncommon among firms, it is more consistent with existing evidence on Italy and other EU countries.[7] Import and export shares exhibit a sizeable degree of variability, with a standard deviation of 5.6 and 26.9 per cent, respectively. In order to gauge the extent of within-firm variability in these external orientation terms, we also calculated the standard deviation of each firm's import and export shares during the sample period. The average values of these dispersion measures, equal to 1.5 per cent for imports and 6.1 per cent for export shares, indicate that most of the sample variability reflects differences across firms. The correlation across firms between imported inputs shares and the shares of foreign over total sales is positive (0.11), and to some extent the low value of this correlation might arguably be a figment of the data limitations in measuring imported input shares at the firm level. In this respect, the empirical literature points to a large correlation across industries between the fraction of importing firms and that of exporting firms, so that internationalized firms tend to be simultaneously exporters and importers (see, for example, Bernard et al., 2007). The mean and median mark-ups are around 9 per cent, with a substantial dispersion. The average and median rates of growth of employment, hours worked, wages and total sales are all rather low, about one per cent in absolute terms. However, in this case there is also a high variability across firms and years, as shown by the high standard deviation, and the negative values of the 25th percentile compared to the positive values of the 75th percentile. Gross job flows at the firm level are sizeable and they also exhibit considerable variability. Average hirings are 10.6 per cent of total employment at the end of the previous year, slightly below average endings (12.1 per cent); the standard deviation, however, is very high (20.4 and 19.8 per cent, respectively) and the median values are sizeably lower than the means. Interestingly, in more than 7 per cent of cases, firms record zero gross hirings.

4 The baseline results

The baseline empirical specification illustrated in section 2 provides a suitable framework for examining the testable implications of our theoretical model. In this section, we will present the

[7] See Bernard et al. (2007) and Mayer and Ottaviano (2008).

results of this specification and of the extensions analyzing the effects of firms' market power on the transmission mechanisms.

4.1 International exposure, employment and hours

The econometric results support the view that exchange rate fluctuations have a significant effect on employment and hours worked. In table 2 (column 1 and 2) we report the findings from estimating equation (2) using the number of employees of each firm as a dependent variable. The difference between the two specifications is that in column 1 the mark-up term is excluded. The coefficient measuring the effect of exchange rate variations through changes in the proceeds from foreign sales is always negative and significantly different from zero at the 95 per cent level (in column 2 it is -1.217 with a standard error of 0.213); the estimated coefficient measuring the effect through the change in costs for imported inputs is always positive and also statistically significant at the 95 per cent level (in column 2 it is 2.681 with a standard error of 1.011). Therefore, an exchange rate depreciation, as measured by a reduction of import and export exchange rates (respectively, $pmer_t$ and $peer_t$) induces an employment expansion through the revenue side of the balance sheet, and a contraction through the cost side. Moreover, the effect stemming from the revenue side increases, in absolute value, with a firm's share of foreign sales in total revenues (χ_{it}), while the effect on the cost side is increasing in the share of expenditure on foreign inputs in total costs (α_{it}).[8]

The signs and magnitudes of the other coefficients estimated from our baseline specification also make sense. Changes in total sales and the level of profit margins (the mark-up) have a positive and statistically significant effect on employment. Similarly, the coefficient of the lagged dependent variable, which accounts for possible persistence in the firm-level adjustment process of the workforce, is positive and statistically significant. The inclusion of a variety of controls in the specification is also justified, as shown by the Wald tests for the joint significance of several groups of dummy variables for time, firm's industry, size and geographic location. Moreover, the value of the Hansen statistic for over-identifying restrictions, testing the hypothesis of lack of correlation between the instruments and residuals, and the value of the test for absence of second-order serial

[8] While, as elucidated in the previous section, we use the permanent component of exchange rate variation in the empirical specification, the results are not significantly altered if actual exchange rate movements are considered. In that case, for example, the estimated response on employment through the revenue side would be -1.017 (with a standard error of 0.194) and the one through the expenditure side would be 2.614 (with a standard error of 0.982). All the results for this and the other specifications of the paper using the actual exchange rate variations, instead of their permanent component, are not reported for space constraints. They are, however, qualitatively very similar to the corresponding findings documented in the paper, and are available from the authors upon request.

correlation of residuals both point to the validity of our specification.[9]

By evaluating the share of export revenues and the share of imported input costs at their median values, the net elasticities of employment with respect to exchange rate variations, estimated through equation (2) (see table 2, column 2), would be 0.05. Hence, assuming identical variations of the export and import exchange rates, the effect of a one per cent depreciation on employment for a hypothetical firm exhibiting these median values of international exposure is almost zero. However, while, on average, the combination of two effects stemming from the revenue and the cost sides point to a negligible estimated elasticity, there may be significant differences in the direction and size of the firm-level response. Perhaps more interestingly, in order to measure the employment sensitivity of firms with different degrees of foreign exposure to exchange rate variations, taking into account the positive correlation between their import and export shares, we estimated the effects on employment for the firms at the 25th, median and 75th percentile of the distribution of the difference between import and export shares, $\alpha - \chi$. For a firm at the 25th percentile (with $\alpha = 0.09$ and $\chi = 0.44$), a one per cent depreciation of the export and import exchange rates determines a 0.29 per cent drop in employment. For a median firm (with $\alpha = 0.12$ and $\chi = 0.22$), employment rises by 0.05. Finally, for a firm at the 75th percentile (with $\alpha = 0.08$ and $\chi = 0.01$), it increases by 0.20 per cent.

While we do believe that our own empirical framework is not ideal for assessing the effect on aggregate employment of exchange rate movements and that other types of data, spanning the whole economy, and other methodologies may be more appropriate for this analysis, we are confident that our approach is particularly suitable for capturing the transmission channels at the firm level. A notable feature of this framework is that it allows us to derive firm-specific, time-varying effects.

The fact that the coefficients associated to $\alpha_{it-1} \cdot \Delta pmer_t$ and $\chi_{it-1} \cdot \Delta peer_t$ have both the expected signs and exhibit large statistical significance is somewhat at odds with the results documented by Campa and Goldberg (2001) for the U.S. economy. In their contribution, when focusing on the whole sample of manufacturing industries, they find that the estimated parameters are not statistically significant, leading them to conclude that the effects of exchange rate swings on jobs and hours worked are relatively small. We believe that this divergence in the findings is not attributable to

[9] In the SIM survey firms' officials were also asked to report their expected value of the number of employees over the subsequent year. We then used expected changes in employment as a dependent variable and ran a specification similar to the one documented in table 2 (column 2). Interestingly enough, also in this case the effect of permanent exchange rate movements on expected employment variation is significant, through both the cost and the revenue side of the balance sheet. In particular, the estimated coefficient associated with $\alpha_{it-1} \cdot \Delta pmer_t$ is 3.058 with a standard error of 0.706, while the one with $\chi_{it-1} \cdot \Delta peer_t$ is -0.911, with a standard error of 0.147 (detailed results are available from the authors upon request).

structural differences between the Italian and the U.S. economy. In this regard, if anything, we would expect that a country like Italy, which has a more rigid labor market, would exhibit a lower net employment elasticity to shocks like an exchange rate swing. On the contrary, our explanation for the more sizable and statistically significant responses that we document deals with the type of data used. Estimations using aggregate industry data are bound to yield an employment elasticity which is lower than the one one obtained with firm-level data. The reason for this lies in the significant employment reallocation patterns across firms characterizing industrialized economies, which have been seen to occur even within narrowly defined industries. Indeed, a large body of literature has focused on this employment (and, in general, resource) reallocation process across plants, pointing to the large heterogeneity at the firm level as the natural explanation. This heterogeneity can be characterized with respect to differences in the development of new products and new production techniques, to differences in entrepreneurial and managerial ability, to the uneven diffusion of information and knowledge, or to the occurrence of firm and plant-specific disturbances. Arguably, aggregating firm-level data does not allow us to capture the within-industry job reallocation driven by currency swings, thus leading to a less significant response of employment to the exchange rate. In order to investigate this issue in more detail, we aggregated our firm-level data into 15 different industries spanning the whole manufacturing sector, and replicated our estimation of equation (2) so that each one of the 182 observations represents an industry rather than a firm. Contrary to our earlier findings, table 3 shows that the estimated effects of $\alpha_{it-1}\Delta e_t$ and $\chi_{it-1}\Delta e_t$ on both employment and hours are not statistically significant when industry-level data are used.

In columns 3 and 4 of table 2 we report the results from estimating equation (2) using the number of hours worked within each firm as a dependent variable, instead of employment. A priori, we would have expected the estimated effect of the exchange rate on hours to be greater (in absolute value) than that on employment, due to the sizeable adjustment costs that typically characterize employment decisions, inducing firms to operate more on the intensive margin (number of hours per employee) and less on the extensive margin (number of employees). Firms' reliance on labor hoarding is particularly important for the Italian economy, where labor market rigidities have been widely documented as significant. During cyclical fluctuations, in general, a firm finds it easier to adjust labor input by using overtime, adding a shift or, during recessions, by resorting to temporary lay-offs (the so called *Cassa Integrazione Guadagni*, which does not affect average employment). Strictly speaking, however, this tendency is more pronounced when the type of shock hitting the producer is transitory rather than permanent. Since we focus on the permanent component of exchange rate movements and, in general, exchange rate shocks tend to be persistent, it is an empirical question as to whether the job response is indeed less significant than that of the hours

worked. Our findings show that the estimated effect of the exchange rate through the cost channel is much more marked on hours than it is on employment (4.775 vs. 2.681, columns 2 and 4). By contrast, when the revenue channel of the balance sheet is considered, the employment response turns out to be stronger than the hour response (-1.217 vs. -0.677). However, if we evaluate the extent of export dependence and import reliance of each firm (χ_{it} and α_{it}) at the median value, it turns out that the elasticity of total manhours with respect to exchange rate variations is equal to 0.45. This value is much larger than the corresponding close-to zero value of the employment elasticity reported before for median values of international exposure. This lends some support to the view that the hours response is, in general, stronger than the employment response.

The coefficients associated with total sale changes and profit margins are positive and statistically significant and, again, the values of the Wald test for the joint significance of the groups of control dummy variables point to their relevance in the equation. Moreover, the values of the Hansen test and of the test statistic for lack of second-order serial correlation of residuals from the first-differenced equation in general lend support to the validity of the specification.

4.2 Market power

As discussed in section 2, the effects of exchange rate fluctuations on labor inputs are shaped by the extent of external orientation both on the cost and revenue side, but they are also amplified for firms with low market power. In table 4 we present the results from estimating our baseline specification, splitting the sample according to the median of firms' mark-ups. As predicted by our model, and consistent with the findings obtained by Campa and Goldberg (2001), the effect of exchange rate fluctuations on employment and hours is stronger for firms with a low mark-up. For these firms, the negative and statistically significant coefficient measuring the effect of exchange rate variations on employment through the revenue channel (-1.069 with a standard error of 0.099) is larger in absolute value than that for high mark-up firms (-0.306 with a standard error of 0.088). Similarly, the coefficient measuring the effect through the cost side is larger for low mark-up firms (2.943 with a standard error of 0.524) than it is for high mark-up firms (1.211 with a standard error of 0.566). As in the previous cases, the test diagnostics confirm the validity of the empirical specification. The response of hours worked to exchange rate changes is also stronger for low mark-up firms. Indeed, as documented in table 4, for these firms the estimated effect on both the cost and revenue channels is, in absolute value, larger than the corresponding effect for firms with relatively high price-cost margins.

In order to investigate more fully how the relationship between exchange rate variations, foreign

exposure and labor inputs is shaped by the extent of firms' market power, we also considered an empirical specification which explicitly includes the market power index described in section 3. In particular, we have used the interaction of the foreign exchange rate variations on the export and import sides with both the external orientation terms and the market power index. In practice, we replaced in equation (2) the two interaction terms on the revenue and cost side with, respectively:

$$(1 - mkup_{it-1}) \cdot \chi_{it-1} \cdot \Delta peer_t \tag{4}$$

and

$$(1 - mkup_{it-1}) \cdot \alpha_{it-1} \cdot \Delta pmer_t \tag{5}$$

This specification resembles more closely the expression for the labor elasticity to the exchange rate, where the components related to both the revenue and cost sides are pre-multiplied by a decreasing function of the mark-up (see, in particular, equation (1)). On the revenue side of the balance sheet, a negative coefficient multiplying the three interacting variables in (4) indicates that an exchange rate depreciation has a positive effect on hours and employment and that this effect increases with a firm's share of export revenues and decreases with its monopoly power. Conversely, on the cost side, a positive coefficient multiplying the interacting terms in (5) implies that an exchange rate depreciation has a negative effect on labor inputs, which increases with a firm's reliance on imported inputs and, again, decreases with its monopoly power.

Table 5 presents the estimation results from running this empirical specification. For both employment and total hours, the coefficients of the two interacting terms written above have the expected sign and are statistically significant. For example, in the equation for total hours, these two coefficients for the revenue and cost channels are, respectively, -0.897 and 3.728 (with a standard error of, respectively, 0.245 and 0.978).

5 Further characterizations

We found that the degree of firms' foreign exposure shapes the labor response to exchange rate shocks. In particular, a high reliance on revenues from foreign sales and on imported inputs magnifies the reaction of jobs and hours worked. We will now consider a number of additional implications of the relationship between exchange rates and employment which, although not all explicitly covered in our simple theoretical framework, can nonetheless find a sound economic rationale and are therefore worth an empirical investigation.

5.1 Input penetration, input substitutability and worker types

An additional transmission channel of exchange rate variations to labor inputs – one that is particularly relevant for firms that are highly exposed to domestic rather than foreign markets – is the degree of import penetration. For a given non-zero value of the exchange rate pass-through elasticity on import prices, a currency depreciation makes imported products more expensive, and therefore less competitive in the internal market. Domestic firms therefore experience a rise in their marginal profitability (and consequently on employment and hours worked), that is larger the higher the share of domestic sales on total revenue $(1-\chi_{it})$, and the higher the degree of competition from foreign suppliers, which can be proxied by the extent of industry's j import penetration (IP_{jt}, measured as the ratio of total import of goods of the j-th industry in domestic demand for goods of the j-th industry).

To see that this prediction is actually contemplated in our theoretical framework, let us consider the expression derived in section 2 for the labor elasticity to exchange rate change (see eq. (1)). In that expression the following entity can be singled out: $A \cdot (1-\chi) \cdot \eta_{p,e}$, where $A = \frac{1}{\bar{\mu}\beta} \frac{a_1}{1+a_1}$ and $\eta_{p,e}$ is the domestic pass-through elasticity measuring the domestic price reaction to exchange rate changes. Under monopolistic competition, the higher the foreign pressure exerted on the domestic market through import penetration, the higher the domestic price elasticity to exchange rate (Dornbusch, 1987). Hence, as stressed by Campa and Goldberg (2001), the domestic price elasticity to the exchange rate is proportional to import penetration, $\eta_{p,e} \propto IP_{jt}$.

To investigate this additional transmission channel of exchange rate movements, we augmented our baseline specification of equation (2) by adding the term:

$$\sum_{j=1}^{K} [\gamma_j (1-\chi_{it-1}) IP_{jt-1} \Delta e_t D_j] \tag{6}$$

where D_j is a dummy variable for each industry j, taking the value of one if firm i belongs to industry j and zero otherwise, and the sector j import penetration ratio, IP_{jt-1}, enters the specification with a one year lag in order to avoid the possible simultaneity effects running from exchange rate to import pressure. The expected sign for the γ_j industry-specific coefficients (with $j = 1, 2, ..., K$) is negative, because an exchange rate appreciation has a stronger negative effect on employment and hours worked for firms that are more dependent on domestic sales in sectors with higher import penetration. Moreover, it is of additional interest to test whether these sectoral coefficients differ among each other.

In table 6 we report the estimation results from running the above specification. The coefficients of

the key variables of the baseline specification continue to have the expected sign and are statistically significant. Interestingly, the Wald test for the joint hypothesis that industry-specific coefficients summarizing the effect of the exchange rate through import penetration are identical is strongly rejected (with a p-value of 0.00), meaning that the extent of import penetration in each industry, combined with differences across firms in the exposure to domestic revenues, introduce another significant source of heterogeneity across sectors and firms in the labor response to the exchange rate. Moreover, in table 7 we report the values of the industry-specific job elasticities to exchange rate changes through the import penetration channel – computed using the industry-specific average value across firms and over time of $(1 - \chi_{it-1})$ and the time average of IP_{jt} for each industry – and the rank for each of these industry estimated elasticities. In the same table, the estimated elasticities can be compared with the corresponding time-average values of import penetration, also accompanied by their ranks. Interestingly, the effects of currency swings on employment through the domestic revenue side are stronger for industries such as computers and office equipment and machinery for industry and agriculture, where the degree of import penetration, and therefore foreign competition, is higher. To investigate the link between these industry-specific revenue-side elasticities and the degree of import penetration, we also computed the Spearman's rank correlation. The value of this coefficient is positive and equal to 0.49, providing further evidence that the labor response to exchange rate induced by foreign pressure in the domestic market is increasing in the degree of import penetration in each industry.

An additional issue is whether the labor response to exchange rate variations through the cost side depends on the degree of substitutability between inputs within the production function. Indeed, if an exchange rate depreciation makes imported inputs more expensive, the effect on marginal profitability and thereby on input choice through the cost side of the balance sheet should depend on how imported inputs can be substituted with domestic inputs. As this depends on technological features of the production process, it is reasonable to expect that it will vary from one industry to the other. In order to tackle this aspect, in an unreported regression we estimated the baseline specification (equation (2)) augmented with sector-specific coefficients associated with the term $\alpha_{it-1}\Delta e_t$. If differences across industries in the degree of input substitutability are indeed relevant, this should introduce another degree of specificity in the total hours response to the exchange rate. We conducted a Wald test for the joint hypothesis that sectoral differences among the coefficients summarizing the cost-side effects on hours of exchange rate changes are equal to zero. The results confirmed that sectoral differences in the hours response through the cost side are indeed statistically significant. However, the effect of an exchange rate appreciation on employment and hours worked remains positive through the cost side and negative on the revenue side for all sectors, consistent with the predictions of our theoretical framework.

Finally, we analyzed whether the reaction of employment and hours worked to exchange rate changes depends on the type of workers employed in each firm. In particular, following Campa and Goldberg (2001), we regressed the estimated firm- and time-specific elasticities of employment and hours worked (obtained, respectively, as $2.681 \cdot \alpha_{it-1} - 1.217 \cdot \chi_{it-1}$ and $4.775 \cdot \alpha_{it-1} - 0.677 \cdot \chi_{it-1}$) on a firm's share of blue- and white-collar workers. Consistent with the fact that, during our sample period, labor market flexibility was lower for blue-collar, table 8 shows that the higher a firm's reliance on blue-collars, the lower the job and hours response to currency swings.

5.2 The firms's export status and the employment-exchange rate link

A large body of empirical evidence in international trade documents that exporting firms are profoundly different from non-exporting firms with regard to a variety of characteristics. The status of exporter for a firm is found to be associated with higher levels of productivity, larger size and higher capital and skill intensity with respect to non-exporting firms (see, for example, Bernard and Jensen, 1999, and Bernard et al., 2007). The substantial microeconomic heterogeneity within narrowly defined industries in a firm's decision to export has led to the development of theoretical models highlighting the interaction of firm-specific features and the export orientation of the firm. The model developed by Melitz (2003), for example, predicts a productivity advantage of exporting firms in light of the existence of a sunk entry cost that induces a self-selection into export markets of the most productive firms. Hence, a higher exposure to trade in the economy leads more productive firms to export whilst, simultaneously, the least productive firms are driven out of the market. The ensuing resource reallocation towards more productive exporting firms is found to account for a significant fraction of aggregate productivity growth (Bernard et al., 2006). According to a different view, a firm's productivity advantage arises instead after its exporting activity has begun and is thus the outcome of a "learning by exporting"process. Whilst the two explanations are not mutually exclusive, the empirical literature, almost unanimously, lends strong support to the former and provides little evidence that entry into export markets induces productivity gains (see Bernard et al., 2007).

In addition to the superior performance features of exporters relative to non-exporters, the empirical evidence on firms' heterogeneity in international trade highlights that the extensive margin of export, that is the number of firms involved in exporting activity, is relatively thin. This is especially true for the U.S., where, as documented by Bernard et al. (2007), the fraction of manufacturing firms that export was only 18 per cent in 2002 (27 per cent in 1997), and for the UK, where it was 28 per cent in 2003 (Mayer and Ottaviano, 2008). The existing evidence points to a larger margin of export in some European countries; for example, Mayer and Ottaviano (2008) report that in

2003 the shares of exporting firms in France, Germany and Italy were, respectively, 67, 59 and 74 per cent, and somewhat lower in Hungary (48 per cent) and Norway (39). Similarly, Castellani et al. (2009) document a participation rate to export markets of Italian firms of about 71 per cent. Whilst some of these differences may be attributable to restrictions in the samples used that often include relatively large firms only, in general, countries exhibit some differences in the degree of internationalization of their firms. Another fact characterizing the pattern of firms' external orientation is the high concentration of aggregate exports among the exporting firms, as relatively few exporters account for the bulk of a country's exports. Trade concentration is documented for the U.S. and for a variety of European countries including, albeit to a lesser extent, Italy (Mayer and Ottaviano, 2008 and Castellani et al., 2009).

As we documented in the data section, our sample covers only manufacturing firms with at least 50 employees and for this reason it is perhaps not suitable for a precise assessment of how uncommon firms' exporting activity is among Italian firms.[10] On the contrary, given our focus on the exchange rate sensitivity of firms' employment and given the evidence on the different characteristics of exporting vs. non exporting firms, we asked ourselves whether differences in the firms' export status introduces a degree of specificity in the employment response to an exchange rate shock. In order to address this issue, we distinguished four alternative cases: 1) firms that are currently exporting; 2) firms that switch export status from non-exporter to exporter; 3) firms that switch status from exporter to non-exporter; and 4) firms that are non-exporters. Therefore not only did we consider how exchange rate affects employment differently for exporters and non-exporters, but we also focused on the employment response around export entry and exit episodes. In panel A of table 9 (column 1) we present the results from the baseline specification estimated on the sample including exporting firms only. Not surprisingly, given firms' high participation rate in the export market in our sample, the results are broadly similar to those reported in table 2, referring to the whole sample. In column 2 we focus on firms in the period after their decision to become exporters and consider a window of three years since exporting has commenced.[11] Our results indicate that exchange rate shocks affect employment significantly through a firm's exposure to foreign product markets but not through the import cost channel. The size of the effect through foreign revenues, however, is smaller than the one estimated in the sample of all exporters. To gauge this, it is

[10] Although largely due to the restricted sample, our data indicates a considerable external orientation of Italian firms along the extensive margin of export, consistent with previous studies. Indeed, by taking the distribution across firms of the share of revenues from export over total revenues, χ_{it}, the export intensity of the firm at the 25th percentile is still different from zero, albeit negligible (about 5 per cent).

[11] Of course, an assessment of the appropriate length of this window is not straightforward and the literature offers little guidance in this regard. For robustness, we therefore replicated the empirical investigation with windows of two and one periods only and the findings are qualitatively unchanged.

important to point out that in the sample of firms that have recently switched export status, not only the estimated coefficient is lower than the one obtained for the generality of exporters (-0.522 versus -1.091), but also the average share of revenues from export over total revenues is sizeably lower (8 versus 30 per cent). Arguably, these results suggest that after taking the important decision to enter export markets, with possible modification in the size of their labor force, firms are less likely to react to exchange rate swings through a further employment adjustment. This is consistent with a large body of evidence suggesting that, whilst higher productivity precedes the decision to undertake exporting activity, it is *after* the entry into the export market that employment growth becomes faster for exporting firms (see Bernard et al., 2007).

For firms switching trading status from exporters to non-exporters, we include in the sample a window of three periods since the firm has exited the export markets. For these firms, and for the non-exporters in general, the implications of currency swings for employment arise in our empirical framework primarily through the imported input channel. This effect, however, is not statistically significant, a result similar to that obtained in the sample of non exporting firms (panel A of table 9, columns 3 and 4). Additional insights can be derived by analyzing the reaction of hours to exchange rate movements. As shown in panel B of table 9 (column 2), after the firm switches status and becomes an exporter, hours are more sensitive than employment to exchange rate changes, and this holds true along both the revenue and the cost channel of the balance sheet. However, given the lower degree of external exposure of firms which have commenced exporting activity, the estimated effect on hours of currency swings is in general lower than the one estimated in the sample of all exporters, consistent with previous findings for employment. On the other hand, a statistically significant hours response to exchange rate variations is uncovered through the import cost side for both non-exporting firms and firms which have lately exited the export market (columns 3 and 4).[12]

[12]It would be interesting to focus also on firms' import behavior and investigate whether the different trading status of importer and non-importer as well as the episodes of entry into and exit from import markets are associated with differences in the exchange rate sensitivity of labor input. The literature on the role of firm heterogeneity in international trade has focused almost exclusively on exporting activity and few studies in this research area have analyzed firms' imports (see, for example, Bernard et al., 2007 and Kugler and Verhoogen, 2009). As we discussed in the data section, however, our measure of reliance on imported input admittedly has some drawbacks and is a substitute for the one we would ideally like to have used. For this reason, we decided to focus on firms' export status only.

5.3 The implications for wages of exchange rate shocks

Firms can react to exchange rate swings not only through changes in their use of labor inputs, but also by modifying the wage that they offer. Based on our theoretical framework (see the appendix), we can postulate the following dynamic empirical specification for wages:

$$\Delta w_{it} = \delta_0 + \delta_1 \alpha_{it-1} \Delta e_{it} + \delta_2 \chi_{it-1} \Delta e_{it} + \delta_3 \alpha_{it-1} + \delta_4 \chi_{it-1} + \\ + \delta_5 \Delta s_{it-1} + \Delta_6 \text{MKUP}_{it-1} + \delta_7 \Delta l_{it-1} + d' Z_{it} + \tau_i + v_{it}, \quad (7)$$

where, as before, lower-case letters refer to the logarithmic transformation of the variable, W_{it} represents real wages per employee, and L_{it} is the employment level of firm i at time t.

Table 10 reports the results from estimating the above specification. Consistent with the model's implications, we find that an exchange rate depreciation induces a real wage increase through the revenue channel. Indeed, the estimated coefficient associated with $\chi_{it-1} \cdot \Delta peer_t$ is -0.722 and it is statistically different from zero (the standard error is 0.140). Moreover, we also document that a depreciation causes the real wage to diminish through the transmission channel based on costs, with an estimated coefficient for $\alpha_{it-1} \cdot \Delta pmer_t$ of 2.608 and a standard error of 0.753. The values of the diagnostic tests provide evidence in favor of the validity of our specification.

These results are in line with those documented by Campa and Goldberg (2001) on industry data. We have argued above that the intense job reallocation process within industries is likely to explain the low employment elasticity to exchange rate changes documented by Campa and Goldberg as compared to our own estimates obtained on firm-level data. The fact that this divergence is significantly reduced when dealing with the estimated wage response to exchange rate changes does not contradict the interpretation of the results for job elasticity and, more generally, it does not diminish the relevance of the job reallocation process, even within narrowly defined industries. Indeed, our findings are consistent with the evidence that the intense job reallocation taking place within industries is usually accompanied by a wage reduction for those workers who are re-employed (see Kletzer, 2000).

Also in the case of wages, the effect of the exchange rate fluctuations is stronger for firms with lower monopoly power. The results presented in the first two columns of table 11, obtained by splitting the sample between low- and high-monopoly power firms, confirm that the wage response to exchange rate changes through the cost channel is stronger for firms with low price-cost margins. The results of the last column of table 11, obtained by introducing in the basic specification the interaction terms described in expressions (4) and (5), also confirm that the sensitivity of wages to exchange rate fluctuations is a decreasing function of a firm's market power on both the cost and

the revenue sides.

5.4 Exchange rate and firm-level gross job flows

Although the focus of this paper is on the effects of exchange rate movements on employment and hours worked, we have also tested if currency swings affect the volume of job creation and job destruction. An important body of literature has focused on the impact of exchange rate fluctuations on inter-industry and intra-industry job reallocation. These contributions appraise the extent of gross and net job reallocations following real exchange movements, and go on to assess how variations in job creation and job destruction bring about changes in net employment. For example, Gourinchas (1998) shows that in the U.S. manufacturing sector, following an appreciation, both job creation and job destruction increases, with the latter rising by more; thus, net employment declines and job reallocation rises (a "churn"). By contrast, following a depreciation, job creation and job destruction decline to the same extent, thus reducing the extent of job reallocation (a "chill"). A somewhat different picture emerges from the French economy, however, where not only job creation and job destruction are more responsive to exchange rate changes than they are in the U.S., but they also move in the opposite direction, with job creation reacting more strongly to the exchange rate than job destruction (Gourinchas, 1999). Interestingly, Klein et al. (2002) also show for U.S. manufacturing that only movements in real exchange rate trends have significant effects for job creation and job destruction, which both increase (decrease) in response to an appreciation (depreciation). Job reallocation is thus significantly affected by a currency movement whilst net employment is not; on the other hand, cyclical variations of real exchange rates only impact on job destruction, with a significant effect on net employment.

In this paper we adopt a narrower perspective than that of the existing studies in this area of research and avoid investigating the degree of inter-industry and intra-industry job reallocation. Instead, we provide a further characterization at the firm level of the employment response after a currency movement. In particular, we use our survey information on job creation and job destruction to investigate whether a given change in net employment induced by exchange rate shocks is associated with high or low levels of job reallocation within the firm.

We use two measures of gross job reallocation as dependent variables: the gross job reallocation rate,

$$gjrr_{it} = \frac{C_{it} + D_{it}}{\frac{1}{2}(L_{it} + L_{it-1})}, \tag{8}$$

and the excess job reallocation rate (Davis and Haltiwanger, 1999),

$$ejrr_{it} = \frac{C_{it} + D_{it} - |C_{it} - D_{it}|}{\frac{1}{2}(L_{it} + L_{it-1})}, \tag{9}$$

where C_{it} and D_{it} are job creation and job destruction, respectively. While the first measure accounts for total job reallocations, the second focuses only on the amount of job reallocation in excess to what is associated with changes in net employment. Table 12 documents the empirical findings from our investigation. The results in the first two columns of table 12 seem to suggest that exchange rate movements induce a significant job reallocation within each firm. Indeed, the coefficients on both the cost and revenue transmission channels are statistically significant with the exception of the effect on $gjrr_{it}$ through the import channel. In the specification where the excess job reallocation rate is the dependent variable, the coefficient associated with $\alpha_{it-1} \cdot \Delta pmer_t$ is 5.137 (with a standard error of 2.453) and the one associated with $\chi_{it-1} \cdot \Delta peer_t$ is -1.457 (with a standard error of 0.472). However, when we turn to investigate separately the job creation rate and the job destruction rate (columns 3 and 4), we find support for the view that exchange rate movements induce significant changes on the rate of job creation, with both coefficients on the revenue and on the cost side having the expected signs and statistically different from zero at the 5 per cent level, but their effects on job destruction are not significantly different from zero. While these results can in no way be compared with those of the previous literature, because our empirical framework is profoundly different, they show that the degree of job reallocation following a currency shock is primarily the result of movements in job creation, rather than in job destruction.

For completeness, we replicated our investigation on "industry-level"data obtained by aggregating our firm-level data up so that an observation in a given year refers to industry. Unlike what we found when we focused on the effect of exchange rate variations on industry-level net changes in employment and hours, we do find a significant effect of the exchange rate on our measures of job reallocation at the industry level. As table 13 shows, the exchange rate significantly affects the extent of sectoral job reallocation. In particular, the effect of a currency appreciation on job reallocation is positive through the imported input channel and negative through the export channel (see columns 1 and 2). The same holds true for the job creation rate at the industry level. For example, the estimated response of the latter variable through the export channel to an exchange rate appreciation is -2.326 with a standard error of 0.962. Interestingly, the extent of job destruction at the industry level also seems to react to exchange rate swings, although to a lesser extent than job creation and job reallocation. In particular, a currency appreciation appears to lower the job destruction rate through the imported input channel in a statistically significant way (the estimated coefficient is -1.845 with a standard error of 0.993), while its effect on job destruction through the export channel is positive, but weakly significant.

6 Conclusion

We document a statistically significant effect of exchange rate variations on employment, hours worked and wages in a representative panel of Italian manufacturing firms. Our primary contribution lies in the investigation of a number of mechanisms at the firm level underlying this finding. We show that the degree of firms' foreign exposure on both the revenue and the cost side of the balance sheet determines the direction and size of the hours and job response to a currency swing. We estimate a time-varying response for each individual firm, reflecting the type and extent of this external orientation. Moreover, we show that, for a given degree of international exposure, the effects of exchange rate variations on employment and hours are magnified when the firm exhibits a low level of monopoly power, and that the response is stronger when firms face a significant foreign pressure on their domestic markets, through a high degree of import penetration. Other features, like the degree of substitutability in the production process between imported and other inputs or the type of workers employed in each firm are shown to introduce an additional degree of specificity in the employment reaction to the exchange rate. We also distinguish four cases to characterize different firms' export status and document that currency swings affect employment differently for exporting and non-exporting firms and also around export entry and exit episodes. This provides another dimension through which firms that export are different from those that do not. We also estimate the wage response to exchange rate shocks, which is significant and again dependent upon firms' foreign exposure. A further characterization of the implications of the exchange rate on labor market variables deals with the amount of gross job flows within each firm following a currency shock. While we provide some evidence that job reallocation within firms is sensitive to exchange rate movements, our findings, however, support the view that job creation is significantly more reactive than job destruction.

Overall, our results show that the aggregate relationship between the exchange rate and the level of employment hinges upon a wide set of characteristics of the firms operating in the manufacturing sector, and therefore currency fluctuations can have very heterogenous effects at the firm level.

Appendix - A simple theoretical model

Let us consider the static problem of a firm that maximizes its profits, π, taking as given: a) the available technology, described by a constant return to scale production function, $Q = F(L, z, z^*)$, where L is labor input and z and z^* are the levels of domestically produced and imported non-labor inputs, respectively; b) the prices of the domestically produced and the imported inputs, respectively s and s^*, expressed in local currency; c) the level of the exchange rate, e (expressed as the number of foreign currency units per unit of domestic currency).

A firm's choice variables are: a) the amount of product to supply in the domestic market and in the foreign markets, respectively q and q^*; b) the volume of domestic and foreign non-labor inputs, respectively z and z^*; c) the amount of labor to be hired, L.

Substituting the inverse demand functions, $p(q, e)$ and $p^*(q^*, e)$ into the profit function, a firm's problem can be defined as:

$$\pi(e) = \max_{q,q^*,z,z^*,L} p(q,e)q + \frac{p^*(q^*,e)q^*}{e} - zs - \frac{z^*s^*(e)}{e} - wL, \tag{A1}$$

subject to the technology constraint: $q + q^* = Q = F(L, z, z^*)$.

The optimal conditions for the solution of this constrained maximization problem are:

$$\frac{\partial p(q,e)}{\partial q}q + p - \lambda = 0, \tag{A2}$$

$$\frac{\partial p^*(q^*,e)}{\partial q^*}\frac{q^*}{e} + \frac{p^*}{e} - \lambda = 0, \tag{A3}$$

$$-s + \lambda\frac{\partial F(L,z,z^*)}{\partial z} = 0, \tag{A4}$$

$$-\frac{s^*(e)}{e} + \lambda\frac{\partial F(L,z,z^*)}{\partial z^*} = 0, \tag{A5}$$

$$-w + \lambda\frac{\partial F(L,z,z^*)}{\partial L} = 0, \tag{A6}$$

where λ is the Lagrange multiplier for the technology constraint and, given (A2) and (A3), it can be expressed as: $\lambda = p(1 + \frac{1}{\eta}) = \frac{p^*}{e}(1 + \frac{1}{\eta^*})$, where η and η^* are the price elasticities of demand in the domestic and foreign product market, respectively. Combining the previous equations and applying Euler's theorem to express total output as the sum of the products of marginal productivity of each factor and the corresponding quantity of input used, simple algebraic manipulations yield the following expression for the logarithm of the optimal demand for labor:

$$\ln L = \ln[\frac{p(q,e)q}{\mu} + \frac{ep^*(q^*,e)q^*}{\mu^*} - (sz + \frac{s^*(e)z^*}{e})] - \ln w, \tag{A7}$$

where $\mu \equiv \frac{\eta}{1+\eta}$ and $\mu^* \equiv \frac{\eta^*}{1+\eta^*}$ are the mark-ups set, respectively, in the domestic and foreign product markets. If we use a standard labor supply equation (e.g., $\ln L = a_0 + a_1 \ln w + a_2 \ln y$, where y is a measure of aggregate demand) to replace $\ln w$ in (A7) and differentiate the resulting equation with respect to the exchange rate, some algebraic manipulations yield the following expression for the elasticity of equilibrium employment, $\widetilde{L}$, with respect to the exchange rate, e:

$$\frac{\Delta \widetilde{L}}{\Delta e} \frac{e}{\widetilde{L}} = \frac{1}{\overline{\mu}\beta}[(1-\chi)\eta_{p,e} - \chi(1-\eta_{p^*,e}) + \alpha(1-\eta_{s^*,e})]\frac{a_1}{1+a_1}, \tag{A8}$$

which is equation (1) in the main text.

In light of equations (A7) and (A8), and following Campa and Goldberg (2001), equilibrium employment can be expressed in reduced form as:

$$\ln \widetilde{L} = b_0 + b_1 \ln e + b_2 \ln y, \tag{A9}$$

where b_1 is equal to the right-hand side of (A8) and b_2 is equal to $\frac{a_2}{1+a_1}$. As discussed in the text, our simple model can be seen as nested in a more general dynamic, stochastic firm's problem with labor adjustment costs that, for our purposes, would deliver the same predictions as those presented above. This hinges on the assumption that the exchange rate follows a random walk, dictating that the best predictor of all its future, uncertain, values is the current value. In this case, as it is shown by Campa and Goldberg (2001) and Nickell (1986), after an exchange rate shock, employment follows a partial adjustment path:

$$\ln L_t = \vartheta \ln L_{t-1} + (1-\vartheta) \ln \widetilde{L}_t, \tag{A10}$$

i.e., each period's level of employment depends on the last period's level and on the current equilibrium level in the absence of adjustment costs; the parameter ϑ, which is a function of the adjustment costs, describes the speed of adjustment to the optimal level. Inserting equation (A9) for $\ln \widetilde{L}_t$ in (A10) yields the following expression for (the log of) employment, which, together with (A7), forms the base of our empirical specification:

$$\ln L_t = c_0 + c_1 \ln L_{t-1} + c_2 \ln e_t + c_3 \ln y_t. \tag{A11}$$

where $c_0 = (1-\vartheta)b_0$, $c_1 = \vartheta$, $c_2 = (1-\vartheta)b_1$ and $c_3 = (1-\vartheta)b_2$.

With regard to wages, by substituting the labor supply equation into (A7) and following a procedure similar to that described above, we obtain the expression for the exchange rate elasticity of wages that provides theoretical underpinnings for equation (7) in the text:

$$\frac{\Delta w}{\Delta e} \frac{e}{w} = \frac{1}{\overline{\mu}\beta}[(1-\chi)\eta_{p,e} - \chi(1-\eta_{p^*,e}) + \alpha(1-\eta_{s^*,e})]\frac{1}{1+a_1}. \tag{A12}$$

Figure 1

Real import and export exchange rates (1998 = 100)

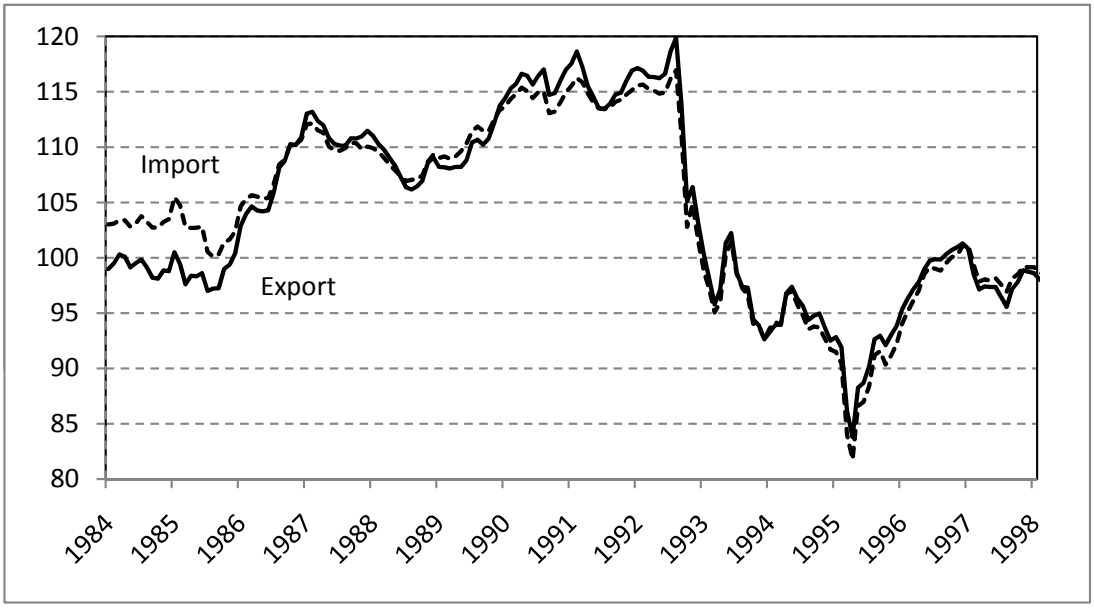

Table 1 Summary statistics

Variable	Mean	St. dev.	Median	25th perc.	75th perc.
Import exchange rate (growth)	−0.29	5.40	0.17	−2.29	3.07
Export exchange rate (growth)	−0.09	5.75	−0.25	−2.72	4.36
Imported input share (α)	13.90	5.62	12.78	10.23	16.40
Export share (χ)	29.77	26.93	23.65	4.89	48.90
Firm-level variability of imported input share	1.45	0.91	1.31	0.74	2.01
Firm-level variability of export share	6.12	6.11	4.45	1.84	8.41
Firm-level variability of import-side interaction	0.74	0.45	0.71	0.37	0.97
Firm-level variability of export-side interaction	1.65	1.60	1.19	0.31	2.63
Mark-up	8.91	25.99	9.34	5.60	13.83
Employment	−1.23	11.31	−1.12	−4.96	2.68
Total hours (growth)	−0.61	16.24	−0.37	−5.99	5.02
Hourly wage (growth)	1.18	16.96	0.90	−2.96	4.82
Average wage (growth)	0.37	18.61	0.53	−3.71	4.78
Gross hirings (share of the labor force)	10.59	20.44	6.31	2.34	12.49
Gross ending	12.05	19.82	8.33	5.16	13.27
Total sales (growth)	0.08	23.64	0.34	−8.58	9.30

Notes: All variables are expressed as percentages. Firm-level variability of imported input share and export revenues are calculated from the time series standard deviation of each firm's imported input share and export share. Firm-level variability of the interactions are calculated from the time series standard deviations of the interaction between each share and its respective exchange rate variation. The sample period considered is 1984-1998.

Table 2: The effect of permanent exchange rate change on employment and hours worked

	(1)	(2)	(3)	(4)
Variable	Employment: Δl_{it}		Hours: Δh_{it}	
$\alpha_{it-1} \cdot \Delta pmer_t$	2.929**	2.681**	3.821**	4.775**
	(1.358)	(1.011)	(1.434)	(1.158)
$\chi_{it-1} \cdot \Delta peer_t$	−1.766**	−1.217**	−0.988**	−0.677**
	(0.329)	(0.213)	(0.275)	(0.237)
α_{it-1}	0.403**	0.410**	0.419**	0.502**
	(.053)	(0.047)	(0.065)	(0.059)
χ_{it-1}	0.003	0.006	−0.001	0.001
	(0.005)	(0.004)	(0.006)	(0.006)
Δs_{it-1}	0.033**	0.025**	0.038**	0.299**
	(0.005)	(0.004)	(0.007)	(0.006)
$MKUP_{it-1}$	−−	0.198**	−−	0.163**
		(0.030)		(0.023)
Lagged dependent variable	0.167**	0.156**	0.031**	0.020**
	(0.016)	(0.012)	(0.013)	(0.011)
Constant	−0.071**	−0.085**	−0.049**	−0.069**
	(0.019)	(0.016)	(0.024)	(0.022)
Year dummies	166.1 (0.0)	150.6 (0.0)	262.9 (0.0)	312.4 (0.0)
Industry dummies	98.7 (0.0)	84.0 (0.0)	112.6 (0.0)	153.8 (0.0)
Geography dummies	17.0 (0.0)	22.1 (0.0)	31.7 (0.0)	36.0 (0.0)
Firm-size dummies	19.1 (0.0)	16.8 (0.0)	36.7 (0.0)	34.3 (0.0)
Hansen test of over-identifying restrictions:	129.2 (0.06)	201.7 (0.08)	189.1 (0.27)	250.0 (0.24)
Test for second-order serial correlation	0.69 (0.49)	−0.08 (0.94)	−1.17 (0.24)	−1.4 (0.15)
Number of observations	6,766	6,766	6,390	6,390

Notes: Estimates are obtained using the system GMM dynamic panel estimator. α_{it-1} is the share of cost for imported inputs and χ_{it-1} is the share of revenues from exports. Δpmer$_t$ and Δpeer$_t$ are the (log) changes in the permanent component of, respectively, the import and export exchange rate. Δs$_{it-1}$ is the (log) change of real sales and $MKUP_{it-1}$ is a firm's mark-up. Size dummies refer to these sizes: 50-99, 100-199, 200-499, 500-999, $\geq$1000 employees. Geographic dummies refer to North-West, North-East, Center, South, Islands. For each group of dummies we report the value of Wald test of their joint significance and the associated p-value. Standard errors are reported in parentheses. They are corrected for heteroskedasticity and incorporate the Windmeijer's finite-sample correction for their original downward bias. The instrument set includes lagged values of the dependent variable, changes of sales and the mark-up dated t-2 and earlier. Hansen is a test of over-identifying restrictions asymptotically distributed as a χ^2. We also report the value of the test for second-order autocorrelation of the differenced residuals (the p-values are reported in parenthes). Sample period: 1984-1998. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 3

The effect of exchange rate on employment and hours using industry data

	(1)	(2)
Variable	Employment: Δl_{jt}	Hours: Δh_{jt}
$\alpha_{jt-1} \cdot \Delta pmer_t$	-12.58	-0.959
	(13.75)	(19.18)
$\chi_{jt-1} \cdot \Delta peer_t$	5.686	2.107
	(7.867)	(10.21)
α_{jt-1}	-0.122	-0.180
	(0.208)	(0.381)
χ_{jt-1}	0.244^{*}	0.225
	(0.138)	(0.169)
Δs_{jt-1}	1.057^{**}	-0.813
	(0.499)	(0.885)
$MKUP_{jt-1}$	2.330^{**}	2.020^{**}
	(0.962)	(0.594)
Lagged dependent variable	-1.368^{**}	0.602
	(0.519)	(0.858)
Constant	-0.346^{**}	-0.232^{**}
	(0.122)	(0.093)
Hansen test of over-identifying restrictions:	4.34 (0.99)	6.54 (0.99)
Test for second-order serial correlation	-1.32 (0.19)	0.94 (0.35)
Number of observations	182	182

Notes: see Table 2. The index j refers to a specific industry. Estimates are obtained by using the system GMM dynamic panel estimator. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 4

Exchange rate, employment and hours: the role of market power (I)

	(1)	(2)	(3)	(4)
Variable	Employment: Δl_{it}		Hours: Δh_{it}	
	Degree of market power		Degree of market power	
	Low	High	Low	High
$\alpha_{it-1} \cdot \Delta pmer_t$	2.943**	1.211**	6.444**	3.264**
	(0.524)	(0.566)	(0.888)	(1.635)
$\chi_{it-1} \cdot \Delta peer_t$	−1.069**	−0.306**	−0.641**	−0.386
	(0.099)	(0.088)	(0.162)	(0.386)
α_{it-1}	0.567**	0.349**	0.532**	0.445**
	(0.058)	(0.041)	(0.073)	(0.074)
χ_{it-1}	−0.005	0.004	0.001	0.007
	(0.005)	(0.004)	(0.006)	(0.007)
Δs_{it-1}	0.011**	0.028**	0.037**	0.035**
	(0.002)	(0.004)	(0.004)	(0.010)
$MKUP_{it-1}$	0.238**	0.273**	0.090**	0.244**
	(0.014)	(0.025)	(0.014)	(0.046)
Lagged dependent variable	0.116**	0.155**	−0.010	−0.003
	(0.009)	(0.009)	(0.010)	(0.015)
Constant	−0.062**	−0.104**	−0.041*	−0.092**
	(0.018)	(0.014)	(0.025)	(0.027)
Year dummies	369.0 (0.0)	228.51 (0.0)	407.1 (0.0)	163.1 (0.0)
Industry dummies	127.9 (0.0)	139.4 (0.0)	47.4 (0.0)	116.7 (0.0)
Geography dummies	28.6 (0.0)	13.4 (0.0)	47.7 (0.0)	11.1 (0.0)
Firm-size dummies	49.2 (0.0)	10.5 (0.0)	30.1 (0.0)	9.6 (0.05)
Hansen test of over-identifying restrictions:	258.5 (0.63)	284.3 (0.22)	279.2 (0.50)	183.0 (0.05)
Test for second-order serial correlation	−0.54 (0.59)	−0.57 (0.57)	−1.80 (0.07)	−1.36 (0.18)
Number of observations	3,339	3,246	3,126	3,099

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. The sample is split based on the degree of firms' market power. The threshold criterion is the median of firms' mark-up. Variables in lower-case letters denote their logarithmic transformation. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 5

Exchange rate, employment and hours: the role of market power (II)

	(1)	(2)
Variable	Employment: Δl_{it}	Hours: Δh_{it}
$\alpha_{it-1} \cdot \Delta pmer_t \cdot (1 - MKUP_{it-1})$	3.859**	3.728**
	(0.752)	(0.978)
$\chi_{it-1} \cdot \Delta peer_t \cdot (1 - MKUP_{it-1})$	−0.875**	−0.897**
	(0.194)	(0.245)
α_{it-1}	0.397**	0.510**
	(0.043)	(0.058)
χ_{it-1}	0.005	0.001
	(0.004)	(0.006)
Δs_{it-1}	0.023**	0.031**
	(0.003)	(0.005)
$MKUP_{it-1}$	0.185**	0.165**
	(0.027)	(0.023)
Lagged dependent variable	0.152**	0.017
	(0.011)	(0.011)
Constant	−0.079**	−0.081**
	(0.014)	(0.021)
Year dummies	172.3 (0.0)	310.6 (0.0)
Industry dummies	90.5 (0.0)	149.7 (0.0)
Geography dummies	21.2 (0.0)	36.7 (0.0)
Firm-size dummies	15.6 (0.0)	35.4 (0.0)
Hansen test of over-identifying restrictions:	238.6 (0.09)	248.7 (0.26)
Test for second-order serial correlation	−0.36 (0.72)	−1.51 (0.13)
Number of observations	6,765	6,389

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. $mkup_{it}$ is the value of the firm's mark-up. Variables in lower-case letters denote their logarithmic transformation. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 6

Exchange rate and employment: the role of input penetration

Variable	Employment: Δl_{it}
$\alpha_{it-1} \cdot \Delta pmer_t$	5.659**
	(2.637)
$\chi_{it-1} \cdot \Delta peer_t$	−1.848*
	(1.123)
$(1-\chi_{it-1}) \cdot \Delta peer_t \cdot IP_{1t-1} \cdot D_1$, $(1-\chi_{it-1}) \cdot \Delta peer_t \cdot IP_{2t-1} \cdot D_2$,....	Wald test:
...,$(1-\chi_{it-1}) \cdot \Delta peer_t \cdot IP_{Kt-1} \cdot D_K$	50.2 (p-val: 0.00)
α_{it-1}	0.414**
	(0.061)
χ_{it-1}	−0.008
	(0.006)
Δs_{it-1}	0.024**
	(0.005)
$MKUP_{it-1}$	0.296**
	(0.034)
Lagged dependent variable	0.175**
	(0.016)
Constant	−0.091
	(0.069)
Year dummies	73.7 (0.0)
Industry dummies	50.0 (0.0)
Geography dummies	13.2 (0.0)
Firm-size dummies	13.5 (0.0)
Hansen test of over-identifying restrictions:	140.7 (0.40)
Test for second-order serial correlation	1.80 (0.07)
Number of observations	5,328

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. IP_{jt} is the value of import penetration experienced by industry j in the year t. D_j is the j-th industry dummy, taking the value of one if firm i belongs to industry j and zero otherwise. The Wald statistic associated with the variables $(1-\chi_{it-1}) \cdot \Delta peer_t \cdot IP_{jt-1} \cdot D_j$ (j=1,2,..., K) tests for the hypothesis that their coefficients are equal. Variables in lower-case letters denote their logarithmic transformation. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 7

Industry heterogeneity in input penetration and the employment response

Industry	Import penetration		Response through import penetration	
	Value	Rank	Estimate	Rank
Transformation of non metallic minerals	0.10	14	−0.11	9
Chemicals	0.35	3	−0.17	7
Metals	0.05	15	−0.02	8
Machinery for industry and agriculture	0.25	5	−1.09	2
Computers, office equipment, precision instruments	0.58	1	−1.51	1
Electrical machinery	0.29	4	−0.22	6
Automobiles and other transport equipment	0.54	2	−0.64	4
Food and tobacco products	0.16	7	0.19	11
Textiles	0.16	8	0.40	15
Leather and footwear	0.17	6	−0.44	5
Clothing	0.14	12	0.28	14
Wood and furniture	0.15	11	0.28	13
Paper and publishing	0.12	13	−0.94	3
Rubber and plastic products	0.15	9	0.14	10
Other manufactures	0.15	10	0.19	12

Notes: Import penetration for each industry is defined in the text and industry time averages are reported here with the corresponding ranks. We also report industry-specific estimates of the exchange rate effect on employment that arises through import penetration. The estimates are obtained from the results of the specification documented in Table 5; we also report the ranks of (the absolute values of) these estimated effects.

Table 8

Workers type and the employment and hours response

Variable	(1) Estimated employment response $2.681 \cdot \alpha_{it-1} - 1.217 \cdot \chi_{it-1}$	(2) Estimated hours response $4.775 \cdot \alpha_{it-1} - 0.677 \cdot \chi_{it-1}$
$\frac{L\_\text{Blue Collars}_{it}}{L_{it}}$	-0.083^{**}	-0.087^{**}
	(0.023)	(0.017)
Constant	-0.266^{**}	0.164^{**}
	(.097)	(.071)
Year dummies	44.4 (0.00)	428.9 (0.00)
Industry dummies	7.8 (0.00)	13.7 (0.00)
Geography dummies	0.36 (0.83)	0.43 (0.78)
Firm-size dummies	5.4 (0.00)	7.79 (0.00)
Hausman test	168.0 (0.00)	130.0 (0.00)
Number of observations	$9,950$	$9,950$

Notes: The dependent variables are the responses of, respectively, employment and hours to permanent exchange rate variations as obtained through the estimation process documented in Table 2. The explanatory variable is the firm-level share of blue-collar workers over total workers. A fixed effects panel estimation method has been applied and values of the Hausman test are reported (with the associated p-value). ** denotes significance at the 5% confidence level and * at the 10% level.

Table 9

(A) The firm's export status and the exchange rate sensitivity of employment

Dependent variable: Employment, Δl_{it}	(1)	(2)	(3)	(4)
Variables	Exporters	Switch to Exporter	Switch to Non-Exporter	Non-Exporters
$\alpha_{it-1} \cdot \Delta pmer_t$	2.236^{**}	-0.625	-0.697	-0.145
	(0.940)	(0.953)	(1.532)	(0.118)
$\chi_{it-1} \cdot \Delta peer_t$	-1.091^{**}	-0.522^{**}	$--$	$--$
	(0.224)	(0.196)		
α_{it-1}	0.445^{**}	0.796^{**}	0.674^{**}	0.419^{**}
	(0.055)	(0.090)	(0.134)	(0.029)
χ_{it-1}	0.008^{*}	0.113^{**}	$--$	$--$
	(0.004)	(0.017)		
Hansen test of over-identifying restrictions:	204.1 (0.06)	82.0 (0.95)	44.8 (0.99)	150.1 (0.41)
Test for second-order serial correlation	-0.08 (0.94)	-0.20 (0.85)	-0.73 (0.46)	-0.60 (0.55)
Number of observations	6,080	267	183	722

(B) The firm's export status and the exchange rate sensitivity of hours

Dependent variable: Employment, Δh_{it}	(1)	(2)	(3)	(4)
Variables	Exporters	Switch to Exporter	Switch to Non-Exporter	Non Exporters
$\alpha_{it-1} \cdot \Delta pmer_t$	2.793^{**}	6.025^{**}	5.659^{**}	2.115^{**}
	(1.187)	(1.972)	(2.350)	(0.141)
$\chi_{it-1} \cdot \Delta peer_t$	-0.616^{**}	-1.127^{**}	$--$	$--$
	(0.271)	(0.563)		
α_{it-1}	0.468^{**}	1.572^{**}	0.457	0.475^{**}
	(0.060)	(0.246)	(0.343)	(0.037)
χ_{it-1}	0.003	-0.001	$--$	$--$
	(0.005)	(0.028)		
Hansen test of over-identifying restrictions:	262.9 (0.07)	70.2 (0.99)	39.4 (0.99)	153.2 (0.70)
Test for second-order serial correlation	0.08 (0.94)	-0.76 (0.45)	-1.72 (0.09)	-1.04 (0.30)
Number of observations	5,753	245	170	671

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. Variables in lower-case letters denote their logarithmic transformation. As in the baseline specification, regressions also include Δs_{it-1}, the (log) change of real sales, a firm's mark-up, $MKUP_{it-1}$, the lagged dependent variable, the constant and Year, Industry, Size and Geographic dummies variables, although the estimation results are not reported. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 10

The effect of permanent exchange rate change on wages

	(1)	(2)
Variable	Real Wages: Δw_{it}	
$\alpha_{it-1} \cdot \Delta pmer_t$	3.739**	2.608**
	(1.015)	(0.753)
$\chi_{it-1} \cdot \Delta peer_t$	−0.726**	-0.722**
	(0.192)	(0.140)
α_{it-1}	0.221**	0.251**
	(0.044)	(0.036)
χ_{it-1}	−0.004	-0.004
	(0.004)	(0.003)
Δs_{it-1}	−0.005	0.001
	(0.004)	(0.003)
$MKUP_{it-1}$	−−	0.126**
		(0.013)
Δn_{it-1}	0.005	0.046**
	(0.012)	(0.009)
Constant	−0.030**	−0.045**
	(0.010)	(0.010)
Year dummies	511.2 (0.00)	809.4 (0.00)
Industry dummies	62.4 (0.00)	168.9 (0.00)
Geography dummies	10.1 (0.04)	10.0 (0.04)
Firm-size dummies	6.5 (0.17)	12.7 (0.01)
Hansen test of over-identifying restrictions:	191.4 (0.09)	283.5 (0.13)
Test for second-order serial correlation	1.31 (0.19)	1.26 (0.21)
Number of observations	6,580	6,580

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. w_{it} is the firm-level value of real wages per employee. Variables in lower-case letters denote their logarithmic transformation. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 11 Exchange rate and wages: the role of market power

Variable	Real Wages: Δw_{it}		
	(1)	(2)	(3)
	Low Mark-up	High Mark-up	All Sample
$\alpha_{it-1} \cdot \Delta pmer_t$	2.848^{**}	2.743^{**}	$--$
	(0.819)	(0.891)	
$\chi_{it-1} \cdot \Delta peer_t$	-1.281^{**}	-0.375^{**}	$--$
	(0.163)	(0.170)	
$\alpha_{it-1} \cdot \Delta pmer_t \cdot (1 - MKUP_{it-1})$	$--$	$--$	2.461^{**}
			(0.848)
$\chi_{it-1} \cdot \Delta peer_t \cdot (1 - MKUP_{it-1})$	$--$	$--$	-1.100^{**}
			(0.184)
α_{it-1}	0.368^{**}	0.215^{**}	0.248^{**}
	(0.055)	(0.049)	(0.040)
χ_{it-1}	-0.007	-0.001	-0.005
	(0.005)	(0.004)	(0.004)
Δs_{it-1}	0.016^{**}	-0.016^{**}	-0.001
	(0.003)	(0.004)	(0.004)
$MKUP_{it-1}$	0.127^{**}	0.168^{**}	0.141^{**}
	(0.012)	(0.034)	(0.016)
Δn_{it-1}	0.083^{**}	-0.015	0.048^{**}
	(0.010)	(0.014)	(0.011)
Constant	-0.053^{**}	-0.051^{**}	-0.048^{**}
	(0.013)	(0.015)	(0.010)
Year dummies	430.2 (0.00)	489.3 (0.00)	576.2 (0.00)
Industry dummies	76.2 (0.00)	85.1 (0.00)	142.0 (0.00)
Geography dummies	14.0 (0.03)	15.7 (0.00)	11.2 (0.02)
Firm-size dummies	6.7 (0.01)	10.0 (0.04)	9.9 (0.04)
Hansen test of over-identifying restrictions:	260.9 (0.16)	210.5 (0.09)	229.3 (0.09)
Test for second-order serial correlation	1.18 (0.24)	1.82 (0.07)	1.29 (0.20)
Number of observations	3,335	3,245	6,579

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. $mkup_{it}$ is the value of the firm's mark-up. When the sample is split based on the degree of firms' market power, the threshold criterion is the median of firms' mark-up. Variables in lower-case letters denote their logarithmic transformation. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 12

Exchange rate implications for gross job flows at the firm level

Variable	(1) Gross Job Reallocation	(2) Excess Job Reallocation	(3) Job Creation	(4) Job Destruction
$\alpha_{it-1} \cdot \Delta pmer_t$	5.567*	5.137**	3.637**	1.645
	(2.986)	(2.453)	(1.691)	(1.536)
$\chi_{it-1} \cdot \Delta peer_t$	−0.443	−1.457**	−1.141**	0.165
	(0.541)	(0.472)	(0.301)	(0.235)
α_{it-1}	−0.076	0.121	0.102	−0.151**
	(0.252)	(0.202)	(0.090)	(0.078)
χ_{it-1}	−0.014	−0.016	−0.006	−0.006
	(0.029)	(0.024)	(0.011)	(0.011)
Δs_{it-1}	0.011**	0.021**	0.023**	−0.002
	(0.005)	(0.004)	(0.003)	(0.003)
$MKUP_{it-1}$	0.032	0.024	0.090**	−0.064**
	(0.056)	(0.056)	(0.031)	(0.021)
Constant	0.258*	0.164	0.126**	0.211*
	(0.152)	(0.121)	(0.041)	(0.040)
Year dummies	126.6 (0.0)	182.9 (0.0)	297.4 (0.0)	114.5 (0.0)
Industry dummies	34.2 (0.0)	33.9 (0.0)	48.7 (0.0)	58.5 (0.0)
Geography dummies	14.5 (0.0)	23.3 (0.0)	35.7 (0.0)	15.3 (0.01)
Firm-size dummies	4.7 (0.32)	7.3 (0.04)	6.3 (0.18)	8.0 (0.09)
Hansen test of over-identifying restrictions:	39.5 (0.83)	58.4 (0.42)	79.2 (0.06)	64.9 (0.34)
Test for second-order serial correlation	−1.83 (0.07)	−0.41 (0.68)	−0.86 (0.39)	−1.37 (0.17)
Number of observations	6,765	6,765	6,765	6,766

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. The four dependent variables are defined in the text and are expressed as rates. Variables in lower-case letters denote their logarithmic transformation. ** denotes significance at the 5% confidence level and * at the 10% level.

Table 13

Exchange rate implications for gross job flows using industry data

Variable	(1) Gross Job Reallocation	(2) Excess Job Reallocation	(3) Job Creation	(4) Job Destruction
$\alpha_{it-1} \cdot \Delta pmer_t$	13.772**	18.34**	4.775**	−1.845*
	(4.420)	(4.759)	(1.928)	(0.993)
$\chi_{it-1} \cdot \Delta peer_t$	−6.523**	−10.91**	−2.326**	0.722
	(2.461)	(3.468)	(0.962)	(0.467)
α_{it-1}	0.021	1.126	0.065	0.050
	(0.184)	(0.706)	(0.104)	(0.131)
χ_{it-1}	−0.038	−0.101*	0.015	0.036
	(0.048)	(0.056)	(0.037)	(0.048)
Δs_{it-1}	0.055**	0.109**	0.033**	0.006
	(0.021)	(0.036)	(0.011)	(0.014)
$MKUP_{it-1}$	0.154	0.381	0.136	1.130**
	(0.616)	(0.571)	(0.270)	(0.438)
Constant	0.201**	−0.011	0.069**	−0.023
	(0.071)	(0.121)	(0.031)	(0.071)
Hansen test of over-identifying restrictions:	6.13 (0.99)	3.93 (0.99)	8.79 (0.99)	8.53 (0.99)
Test for second-order serial correlation	1.55 (0.12)	1.28 (0.20)	1.04 (0.30)	−0.89 (0.37)
Number of observations	156	156	156	156

Notes: see Table 2. Estimates are obtained by using the system GMM dynamic panel estimator. The four dependent variables are defined in the text and are expressed as rates. Variables in lower-case letters denote their logarithmic transformation. ** denotes significance at the 5% confidence level and * at the 10% level.

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