Exchange Rate Shocks and Trade: Firm-Level Evidence from EU Enlargement

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December 2014

Abstract

This paper examines the role of imported inputs in cushioning exchange rate shocks to exporters. The implications of a partial equilibrium model of heterogeneous firms and trade are tested. Namely, I estimate the impact of a hypothetical 1% appreciation of the domestic currency on sales using a panel of Czech firms. In the model, an exchange rate shock simultaneously affects the variable costs and revenues associated with exports and imported inputs. According to the data, the Czech Republic's EU accession in 2004 boosted firms' participation in export and import. This rich within-firm variation in trade strategies was used to identify the exchange rate elasticities of trade. For firms that both export and import, export sales are predicted to drop only by 0.8% compared to a 1% drop for a price-taker exporter who does not use imported inputs.

JEL Codes: C23, C26, D22, D24, F12.

Keywords: Exchange rate shocks, heterogeneous firms, international trade, productivity, production function.

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Most of this work was carried out while the author was employed at the Czech National Bank. The views expressed in this paper are those of the author and not necessarily those of the National Bank of Slovakia or the Czech National Bank. The author is responsible for all errors and omissions in the present work.

Acknowledgement: The author is grateful to the Czech National Bank for providing support and access to the dataset used in this study. The author would like to thank Randall Filer, Kamil Galuščák, Jan Hagemejer, László Halpern, Jan Kmenta, Lubomír Lízal, Branislav Saxa, Jan Švejnar, Petr Zemčík, and Krešimir Žigić for their helpful comments and suggestions.

1. Introduction

Over recent years in the Czech Republic, we have witnessed anecdotal evidence of domestic currency appreciation causing alarm among heads of large export-oriented industrial companies and industrial associations. These managers argued that a strong domestic currency negatively impacted the profit margins of Czech exporters, as export prices are usually contracted in foreign currency. At the same time, it is a well-known fact that the import intensity of Czech manufacturing exports has been high, especially since the Czech Republic joined the EU. This paper investigates the extent to which cheaper imported intermediate products compensate for a drop in export sales as a result of appreciating local currency. Our answer to this question will be based on a model-backed estimate using firm-level panel data.

We apply a partial equilibrium model with monopolistically competing firms which are heterogeneous in their productivities. In the model setup, firms will serve the domestic market, export final goods, or import inputs, depending on their productivities. Next we introduce an exogenous exchange rate shock, which simultaneously affects the variable costs and revenues associated with exports and imports. This allows us to estimate the impact of a hypothetical 1% appreciation of the domestic currency on sales according to different trade strategies. The predictions above will follow from the equilibrium sales equation implied by the model. The equation relates the log of total sales to exporting, importing and productivity and their coefficients are combinations of the model's structural parameters.

In our effort to identify the coefficients in the sales equation, we face two main econometric problems. The first concerns the fact that firms do not select into exporting and importing strategies randomly. According to the model, the selection is based mainly on the productivity of the firm and other industry-specific parameters. To correct the potential selectivity bias in the coefficients of exporting and importing, we instrument them by the fitted probabilities of firms engaging in those activities. The probabilities are estimated from a year-by-year multinomial probit model. The model considers the choice between serving the domestic market only, exporting in addition, importing in addition or engaging in all these activities. The second problem is represented by the productivity variable, which needs to be estimated. We fit total factor productivity from a standard firm-level production function extended by the possibility of using imported intermediate goods. Following recent studies in the literature, we use GMM and instrumental variable estimation to correct for the measurement error in capital.

To estimate the exchange rate elasticities we use an unbalanced panel of 7,356 Czech manufacturing firms observed from 2003 to 2006. The studied interval is crucial for the identification of the estimates, as it can be characterized by high within-firm variation in exporting and importing strategies. The variation is likely to be associated with the exogenous lifting of trade barriers due to Czech EU membership since 2004. This motivated an increasing share of firms to engage in importing intermediate goods and exporting final products.

The remaining part of the paper is organized as follows. Section 2 reviews the relevant literature, Section 3 sets up the model, Section 4 outlines the testable implications of the model, Section 5 describes the dataset, Section 6 explains the estimation procedure, Section 7 summarizes the results, and the last section concludes.

2. Literature Review

Based on theory and empirical evidence, more productive and larger firms are more likely to import and export products than their less productive and smaller competitors. This is explained by the fixed costs associated with serving foreign markets and maintaining distribution networks, i.e., economies of scale.[2] In addition, recent firm-level evidence suggests that importing intermediate goods tends to improve the productivity of firms. This productivity gain is explained by the higher quality of imported intermediates or the higher degree of differentiation of the final good. In what follows we first summarize papers that have studied the productivity-increasing effect of imports in the context of heterogeneous firms. Second, we briefly outline papers that have considered both exports and imports in the same setup. Third, we mention microeconomic studies that have dealt with exchange rate shocks. Finally, we position our paper in the literature.

First, there are several theoretical and empirical studies that investigate the connection between firm heterogeneity in productivity, importing, and exporting. For example, Kasahara and Rodrigue (2008) find evidence that importing intermediate goods improves plant performance in Chilean manufacturing firms. The authors extend a standard Cobb-Douglas production function with capital, labor, and material inputs to include a binary indicator of importing. While estimating the production function, they address the simultaneity issue of inputs and the productivity shock using a two-stage GMM procedure.

Halpern et al. (2011) use product-level customs data merged with a panel of Hungarian firms. Their findings suggest an increase in firm productivity due to a higher fraction of imported product varieties used. Accordingly, about two-thirds of this productivity gain is estimated to come from greater diversification of inputs and thus a more differentiated final good. The rest of the gain can be attributed to the higher quality of imported intermediates. Finally, this study also estimates the impact of a hypothetical tariff cut on imports and the number of input varieties. The above estimate is available thanks to the identification of some of the model's structural parameters, which also involves fitting a production function. The approach to estimating the production function is similar to that of Kasahara and Rodrigue (2008).

Second, Helpman et al. (2004) introduce a model of heterogeneous firms facing the decision to serve only the domestic market or to additionally access foreign markets by exporting or through horizontal foreign direct investment. Firms in this model sort into various organizational forms according to their productivities. The least productive firms serve the domestic market only. More productive firms serve the domestic market and export to foreign markets at the same time. Firms with the highest productivity set up production plants abroad to serve the foreign market. The authors find support for the above ranking of firms based on industry-level estimates using data on exports and FDI sales of U.S. firms.

Kasahara and Lapham (2013) develop a dynamic model with heterogeneous firms which can opt to import intermediates and export to foreign markets. The authors estimate the structural parameters of their dynamic model using a complex nested likelihood function on a Chilean panel of firms. They also perform counterfactual experiments of policy changes affecting trade barriers, such as tariffs. Their experiments suggest that trade improves aggregate productivity and welfare. Furthermore, policies increasing import barriers can inhibit the export of goods.

[2] The idea of economies of scale in exporting under monopolistic competition dates back to Krugman (1980) with homogeneous firms and Melitz (2003) with heterogeneous firms.

Bas and Strauss-Kahn (2011) use a static model of heterogeneous firms with exports and imports to study the effect of the number of input varieties on TFP and export sales. The authors use a French combined firm- and product-level dataset similar to the Hungarian data of Halpern et al. (2011). In addition, Bas and Strauss-Kahn's 2011 model extends that of Halpern et al. (2011) by considering the possibility of firms exporting. The authors test the model's implications as partial correlations between certain variables of interest, although the estimating equation does not come directly from the model, nor do they estimate structural parameters.

Third, some theoretical papers have dealt with the problem of exchange rate pass-through to domestic prices and firm sales from a microeconomic point of view. For example, Jäger (1999) studied the impact of an exchange rate shock on prices in a two-country duopoly. The two firms are registered in different countries, but each of them serves both markets with a homogeneous final good. Baniak and Philips (1995) study the effect of an exchange rate shock on prices and sales in a two-country duopoly model with the joint production of two commodities by each firm. The authors look at the interaction between the exchange rate shock on the one hand and strategic substitutability and complementarity of goods produced, economies of joint production of two final goods, and economies of scale on the other hand.

The main disadvantage of the duopoly models mentioned above is that they ignore the possibility of differentiated products, firm heterogeneity, and the resulting co-existence of trading and non-trading firms in an industry. The monopolistically competing heterogeneous firms approach is thus closer to what is normally observed in firm-level data. However, the latter approach disregards the possibility of competition from foreign producers and the impact of tariffs or exchange rate shocks through this channel.

Finally, we clarify the connection between the existing literature and our setup. Combining two branches of static models, we consider exportation and importation by monopolistically competing heterogeneous firms in partial equilibrium. First, we use the core of the model by Helpman et al. (2004), including exports, but ignoring the possibility of FDI. Second, we extend this model to include productivity-improving imported intermediates, similarly to Kasahara and Rodrigue (2008). However, due to data limitations, we do not study the effect of input varieties on TFP or exports as in Halpern et al. (2011) and Bas and Strauss-Kahn (2011). Using estimates of the model's equilibrium sales equation we compute the exchange rate elasticities of domestic and export sales for Czech manufacturing firms.

To sum up, the present paper offers a static alternative to Kasahara and Lapham (2013) with the advantage of a simpler model and a computationally less intensive estimation procedure. In contrast to Bas and Strauss-Kahn (2011), we test the implications of the model through the equilibrium sales equation obtained directly from the model.

As perhaps the main novelty of this paper, we study the effect of exchange rate shocks on firm sales. To our knowledge, currency shocks have not been studied in the context of heterogeneous firms and trade. In the related literature it is typical to estimate the more straightforward impact of an import tariff change. In light of the establishment of several free trade areas worldwide in recent decades, tariff changes have become less frequent and also less relevant for current macroeconomic policy compared to exchange rate shocks.

3. The Model

We consider N sectors in the economy, each of which produces differentiated products. Consumer expenditures on each sector's total output are exogenously fixed. At the beginning of a period each firm i in a given sector receives a productivity shock e_i. After e_i is revealed, firms decide whether to do business in their sector or not. If production will take place, firms can choose whether to serve the domestic market only ($X=0$) or additionally to export ($X=1$). Furthermore, firms can decide to use domestic intermediate goods only ($M=0$) or to employ a mix of domestic and imported intermediates ($M=1$). The decisions of firms to export or import will influence their fixed and variable costs associated with trade. Moreover, in the case of production including imported intermediates, firms' productivity will increase to $e_i(M=1) = ne_i > e_i(M=0) = e_i$. As in Kasahara and Rodrigue (2008), we attribute this increase in productivity to higher quality of foreign intermediates or to the variety effect stemming from a more differentiated final good.[3]

Trading decisions are subject to the following fixed and variable costs. Running a production plant necessitates spending a fixed cost f. Serving foreign markets bears additional fixed costs f_X associated with expenditures on marketing and maintaining logistic networks abroad. Similarly, importing intermediates also involves extra fixed costs f_M. Participation in trade is additionally associated with variable costs of transportation. As is common in the literature, we assume melting-iceberg transport costs for exports $\tau_X > 1$ and imports $\tau_M > 1$, which require τ units to be shipped for one unit to arrive. The full structure of variable costs $c(X,M)$ and fixed costs $f(X,M)$ looks as follows:

$c(X=0, M=0) = c,$	$f(X=0, M=0) = f,$
$c(X=0, M=1) = c\tau_M,$	$f(X=0, M=1) = f + f_M,$
$c(X=1, M=0) = c\tau_X,$	$f(X=1, M=0) = f + f_X,$
$c(X=1, M=1) = c\tau_M\tau_X,$	$f(X=1, M=1) = f + f_M + f_X$

Firms compete in monopolistic competition[4] and preferences across varieties within a sector are modelled by a CES utility function.[5,6] The elasticity of substitution between varieties within a sector is a constant $\varepsilon = 1/(1–\alpha) > 1$, where $1/\alpha$ is the monopolistic price mark-up. Monopolistic competition and CES preferences imply the following demand function for the product of firm i in market j:

[3] In the absence of product-level information on imported intermediates matched to firm-level data we are unable to differentiate the two effects empirically. Halpern et al. (2011) study such disaggregated data and conclude that two-thirds of the increase in firm productivity when imported intermediates are used is due to the variety effect.

[4] As monopolistic competition assumes an infinite number of atomistic firms producing different varieties of a good, we checked the degree of market share concentration within each manufacturing sector by two-digit NACE codes. Using the standard Herfindahl index of sales, all sectors were found to be highly unconcentrated, with index values below 0.01. Note that the Herfindahl index ranges from 0 to 1 and is computed as:

$H = \sum_{i=1}^{N}(s_i^2)$, where s_i is the market share of firm i and N is the number of firms.

[5] The CES utility function over h varieties of goods $\mathbf{x}$ within a sector takes the standard form:

$u(\mathbf{x}) = (x_1^{\alpha} + x_2^{\alpha} + \ldots + x_h^{\alpha})^{1/\alpha}$, where $\alpha = (\varepsilon-1)/\varepsilon$

[6] The assumption of CES utility can be relaxed while maintaining the main results of the model. Mrázová and Neary (2011) show that if the operating profits function satisfies supermodularity conditions, the equilibria of the model and the productivity cut-offs in Figure 1 can be maintained. Supermodularity would be satisfied, for example, by quadratic preferences, other things being equal. We leave extensions of the model in this direction for future research.

$q_{ij} = A_j p_{ij}^{-\varepsilon}$ (1)

where A_j is the constant sectoral demand level in market j, with values $A_{j=0} = A$ for the domestic market and $A_{j=x} = A_x$ for the foreign market. The values of A_j are assumed to be exogenous to the firm.

The production function is a simplified version of Kasahara and Rodrigue (2008) and extends Helpman et al. (2004) by introducing productivity-increasing imported intermediates. We define production as:

$q_i = e_i(M)I_i$ (2)

where $e(M)$ is the productivity coefficient as a function of the binary import indicator M, and I_i is the amount of intermediate goods used in production.

Using demand (1), production (2), and cost functions $c(X,M)$ and $f(X,M)$ we can write firm i's profit from serving market j as:

$$\Pi_{ij}(M) = A_j p_{ij}^{1-\varepsilon} - c(X,M)I_{ij} - f(X,M) = A_j p_{ij}^{1-\varepsilon} - c(X,M)q_{ij}/e_i(M) - f(X,M) =$$

$$= A_j p_{ij}^{1-\varepsilon} - c(X,M)A(X)p_{ij}^{-\varepsilon}/e_i(M) - f(X,M) \quad (3)$$

The profit-maximizing unit price then becomes:

$p_{ij}^* = p_i^* = \varepsilon c(X,M)/[e_i(M)(\varepsilon\text{-}1)]$ (4)

Plugging the above equilibrium prices (4) into the profit function (3) we get the following equilibrium profits for various trade strategies:[7]

$\Pi_i^*(X,M) = \Pi_{i0}^*(M) + \Pi_{ix}^*(M)$ (5)

$\Pi_i^*(0,0) = EA\ [e_i(0) / c]^{\varepsilon-1} - f$

$\Pi_i^*(0,1) = EA\ [e_i(1) / c\tau_M]^{\varepsilon-1} - f - f_M$

$\Pi_i^*(1,0) = E(A+A_x\tau_X^{1-\varepsilon})\ [e_i(0) / c]^{\varepsilon-1} - f - f_X$

$\Pi_i^*(1,1) = E(A+A_x\tau_X^{1-\varepsilon})\ [e_i(1) / c\tau_M]^{\varepsilon-1} - f - f_M - f_X$

where $E = \varepsilon^{-\varepsilon} (\varepsilon\text{-}1)^{\varepsilon+1}$ is a positive constant. In equilibrium, each firm i will select the trade strategy (X,M) with the highest profit for firm i or will exit if none of $\Pi_i^*(X,M) > 0$.

Note that all parameters of $\Pi_i^*(X,M)$ are constant for a given sector, except the firm-specific productivities e_i. Thus, the equilibrium trade strategies (X,M) within a sector will differ only by e_i. Plotting all $\Pi_i^*(X,M)$ against $[e_i(0)]^{\varepsilon-1}$ results in a linear graph which offers helpful insights into the model's equilibrium trade strategies (Figure 1). Notably, we find firms in our data selecting into all four (X,M) strategies within each manufacturing subsector.[8] Therefore we focus on a set of parameters that implies the existence of all trade strategies in sectoral equilibrium.

[7] Note that equilibrium requires $\Pi_{ij}^* > 0$.

[8] In our empirical analysis we use the first two digits of the firms' NACE codes. NACE is a European standard for classifying the economic activity of firms.

Furthermore, we assume the following ranking of cut-off productivities that imply equilibrium trade strategies for firms in terms of e_i: $0 < e_{00} < e_{10} < e_{01} < e_{11}$. This means that the least productive firms, with $e_i < e_{00}$, will not do business. Next, firms with e_i falling into any of the latter four intervals will optimally choose the *(X,M)* strategy as indicated by the subscript of each interval's lower bound e_{XM}. The ranking of productivity cut-offs above is justified by our data. As we will show in the Data section below,[9] the average firm size in the sub-samples by trade strategies follows the same order as our assumption about the productivity ranking. In the model, a higher productivity coefficient e_i implies larger profits, revenues, and thus firm size.

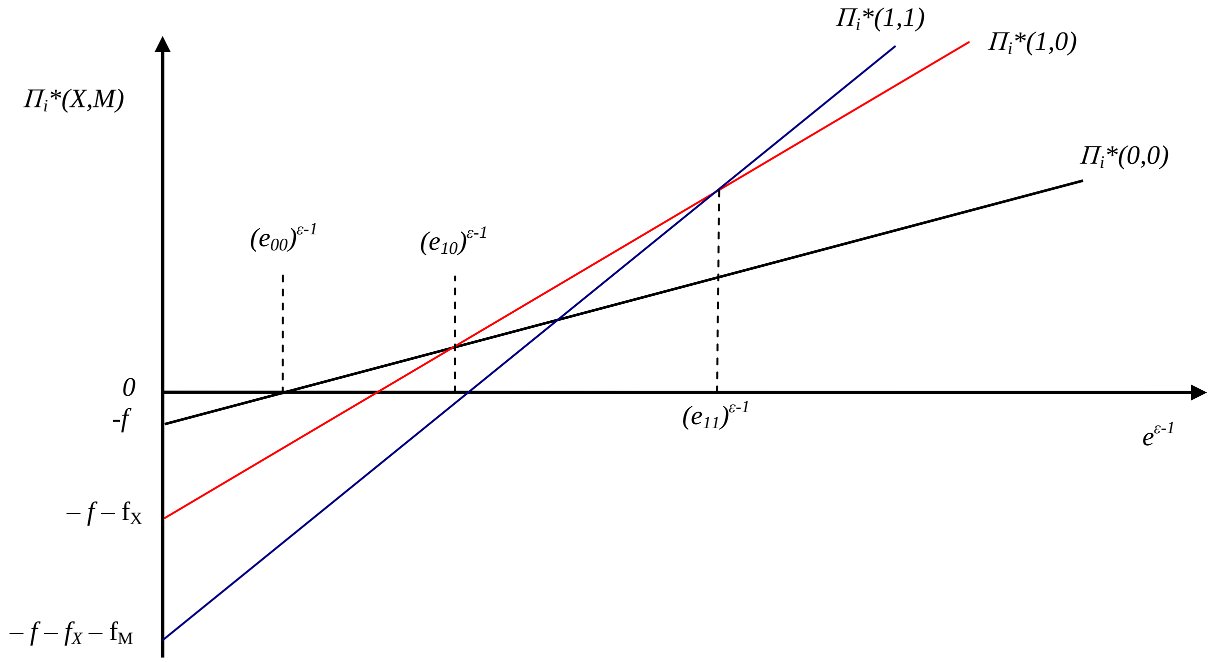
We can argue that if all *(X,M)* strategies are to be observed in sectoral equilibrium, e_{00} must come first and e_{11} last. This is because the slope of $\Pi_i^*(1,1)$ with respect to $[e_i(0)]^{\varepsilon-1}$ is the highest and the intercept the smallest among $\Pi_i^*(X,M)$. The other extreme is $\Pi_i^*(0,0)$, with the smallest slope and the largest intercept. Although both alternative positions of e_{10} and e_{01} can exist in different sectoral equilibria, we will discuss only the $e_{10} < e_{01}$ case as suggested by our data. In the following we outline the assumptions about the parameters of $\Pi_i^*(X,M)$ other than e_i that are necessary to arrive at the productivity ranking mentioned above.

If $\Pi_i^*(0,0)$ is to earn positive profits, productivity e_i must exceed the cut-off point $(e_{00})^{\varepsilon-1} = (fc^{\varepsilon-1}) / EA$. Given that $\Pi_i^*(0,1)$ and $\Pi_i^*(1,0)$ have a lower intercept than $\Pi_i^*(0,0)$, strategies *(0,1)* and *(1,0)* will exist in equilibrium only if the slopes of $\Pi_i^*(0,1)$ and $\Pi_i^*(1,0)$ with respect to $[e_i(0)]^{\varepsilon-1}$ are greater than the slope of $\Pi_i^*(0,0)$. This requires $[n / \tau_M]^{\varepsilon-1} > 1$ in the case of $\Pi_i^*(0,1)$ and $A_x\tau_X^{1-\varepsilon} > 0$ for $\Pi_i^*(1,0)$. From inequalities $e_{10} < e_{01}$, $e_{00} < e_{01}$, and $e_{00} < e_{10}$ we get further conditions. Assuming that $f_M > f_X$ and $A(n/\tau_M)^{\varepsilon-1} > (A+A_x\tau_X^{1-\varepsilon})$ will ensure that the equilibrium is located within the relevant positive range of $[e_i(0)]^{\varepsilon-1}$, where the latter inequality is the relationship between the slopes of $\Pi_i^*(1,0)$ and $\Pi_i^*(1,0)$ with respect to $[e_i(0)]^{\varepsilon-1}$. The condition $e_{10} < e_{01}$ further requires $f_M(A^{-1}A_x\tau_X^{1-\varepsilon}) > f_X[(n/\tau_M)^{\varepsilon-1} - 1]$.

The remaining equilibrium profit function, $\Pi_i^*(1,1)$, has the lowest intercept of all the trade strategies, amounting to $- f - f_M - f_X$. The profit of the strategy of simultaneously exporting and importing will thus exceed that of other strategies if and only if the slope of $\Pi_i^*(1,1)$ with respect to $[e_i(0)]^{\varepsilon-1}$ is larger than the slopes of the other three $\Pi_i^*(.,.)$. This requires $[n / \tau_M]^{\varepsilon-1} > 1$ and $A_x\tau_X^{1-\varepsilon} > 0$, which is in accordance with all the assumptions above. Figure 1 depicts the sectoral equilibrium with profit lines for different trade strategies.

[9] See sales, real value added, real capital, labor, energy, and material inputs in Table 4 in the Data section and Table A1 in Appendix 1.

Figure 1: The Most Productive Firms Import and Export (the least productive entrants do not trade)



Note: For better trackability of the figure let us assume that $\Pi_i*(1,0) = \Pi_i*(0,1)$ and $f_X = f_M$.

4. Testable Implications

In this section we derive the estimable equilibrium sales[10] equations of our model. The estimates from the sales equations enable us to quantify the impact of a hypothetical exchange rate shock on firm sales depending on different trade strategies. At the end of the section, the exchange rate elasticity estimates obtained from the sales equation are derived.

Using (1) and (4), the equilibrium sales equation of firm i serving market j can be written as:

$$S_{ij}(X,M) = A_j(p_{ij}*)^{1-\varepsilon} = A_jE'c(X,M)^{1-\varepsilon}e_i(M)^{\varepsilon-1} \qquad (6)$$

where $E' = [\varepsilon/(\varepsilon\text{-}1)]^{1-\varepsilon}$ is a positive constant. Using (6) we can also write total sales in all markets served as a function of trade strategies:

$$S_i(X,M) = S_{i0}(X,M) + S_{ix}(X,M) \qquad (7)$$

$$S_i(0,0) = AE'c^{1-\varepsilon}e_i(0)^{\varepsilon-1}$$

$$S_i(0,1) = AE'(c\tau_M)^{1-\varepsilon}e_i(1)^{\varepsilon-1}$$

[10] We estimate sales equations rather than equilibrium profits, as in the former case we do not need to identify the fixed cost parameters *f(X,M)* for the exchange rate elasticity estimates. Note that in order to estimate fixed costs we would need further identifying assumptions.

$S_i(1,0) = (A+A_x\tau_X{}^{1-\varepsilon})E'c^{1-\varepsilon}e_i(0)^{\varepsilon-1}$

$S_i(1,1) = (A+A_x\tau_X{}^{1-\varepsilon})E'(c\tau_M)^{1-\varepsilon}e_i(1)^{\varepsilon-1}$

Now let us introduce the exchange rate into the above sales equations with the aim of estimating the impact of a hypothetical exchange rate shock. We assume that the exchange rate $r > 1$ expresses the value of the foreign currency in terms of the domestic currency.[11] Furthermore, connecting to our anecdotal evidence from the Czech Republic mentioned in the introduction, we will study a shock of an appreciating domestic currency reducing r. This results in decreased variable costs of acquiring imported intermediates τ_M and thus higher equilibrium profit and sales. At the same time a stronger domestic currency implies a decreased demand level on export markets A_x measured in the domestic currency. We examine the instant impact of the exchange rate shock on profit and sales assuming that the prices of imported intermediates and exported final goods are contracted in the foreign currency and that the firm is unhedged against currency movements. The next paragraph lends some support to our assumptions above.

Recent survey evidence by Čadek et al. (2011) on the hedging behavior of 184 Czech exporter firms in the period 2005–2009 relates to our assumptions regarding the exchange rate shock. Specifically, more than 75% of the exports of the firms surveyed are contracted in euros and about 90% go to the Eurozone and the rest of Europe. Next, about 30% of the respondents were fully unhedged against currency movements. Furthermore, about 50% of those who at least partially hedge their foreign currency exposure use so-called natural hedging. This involves the temporal alignment of cash inflows and outflows denominated in foreign currencies. As is known, natural hedging does not perfectly eliminate foreign currency risk. Finally, the typical hedging horizon among the respondents was also in line with our assumption of a short-run effect. Specifically, about 80% of the hedgers typically considered a horizon of less than one year.

Now we implement the exchange rate shock in equations (6) and (7). According to the model, firms with different trade strategies are affected differently by the exchange rate shock.[12] Those which do not export and import will not be impacted. Next, firms using imported inputs will be able to offer their product at a lower price and their equilibrium sales will increase, ceteris paribus. Furthermore, firms serving export markets will experience a decrease in their equilibrium export sales due to a lower demand level. Finally, the net effect of the exchange rate shock on the total sales of firms that both export and import can be either positive or negative. This is because their sales on domestic markets will increase due to cheaper imported inputs. At the same time, the negative effect of lower export demand may or may not fully outweigh the positive effect of cheaper imported inputs on export sales.

We can incorporate the exchange rate r into the equilibrium sales equations (7) as follows:

$$S_i(0,1) = S_{i0}(0,1) = AE'[c\tau_M r]^{1-\varepsilon}e_i(1)^{\varepsilon-1} \quad (8)$$

$$S_i(1,0) = S_{i0}(1,0) + S_{ix}(1,0) = (A+rA_x\tau_X{}^{1-\varepsilon})E'c^{1-\varepsilon}e_i(0)^{\varepsilon-1} \quad (9)$$

$$S_i(1,1) = S_{i0}(1,1) + S_{ix}(1,1) = (A+rA_x\tau_X{}^{1-\varepsilon})E'[c\tau_M r]^{1-\varepsilon}e_i(1)^{\varepsilon-1} \quad (10)$$

[11] This is CZK/EUR in the Czech case.

[12] Here we focus on the intensive margin only, which means discussing the partial effects on firms in a given equilibrium trade strategy. At the same time we ignore the extensive margin, i.e., the effect of the exchange rate shock on some firms changing their trade strategies.

The equations above imply the following exchange rate elasticities of sales for the trade strategy *(X,M)* and the market served *j*, where *j=0* denotes the domestic market and *j=x* denotes export markets:

$$\rho_j(X,M) = (r / S_{ij})\, \partial S_{ij} / \partial r$$

$$\rho_0(0,1) = \rho(0,1) = \rho_0(1,1) = (1\text{-}\varepsilon) \qquad (11)$$

$$\rho_x(1,0) = 1$$

$$\rho_x(1,1) = (2 - \varepsilon) \qquad (12)$$

$$\rho(1,1) = [(1-\varepsilon)A + (2-\varepsilon)rA_x\tau_X^{1-\varepsilon}]/(A + rA_x\tau_X^{1-\varepsilon}) =$$

$$= [1 - \varepsilon + rA_x\tau_X^{1-\varepsilon}/(A + rA_x\tau_X^{1-\varepsilon})] =$$

$$= 1 - \varepsilon + R \qquad (13)$$

where ratio $0 < R < 1$ on the right-hand side of the above equation is the share of the freight cost-discounted foreign demand level $rA_x\tau_X^{1-\varepsilon}$ in the total demand level faced by exporters.

Given that the elasticity of substitution between varieties in a given sector ε is assumed[13] to be greater than one, we expect a negative exchange rate elasticity of domestic sales $\rho_0(.,1)$. This means that the shock of an appreciating domestic currency implies positive sales growth on domestic markets for firms that import some of their intermediates. Furthermore, according to the equations above, export sales are unit elastic to the exchange rate when no intermediates are imported and thus will decrease if the home currency appreciates. Next, the elasticity of export sales if some intermediates are imported $\rho_x(1,1)$ is negative if $\varepsilon > 2$ and non-negative if $1 < \varepsilon < 2$. Hence it follows that firms with trade strategy *(1,1)* can still experience increased export sales due to the exchange rate shock, i.e., $\rho_x(1,1) < 0$, if ε is large enough. In the above case the positive effect of cheaper imported intermediates outweighs the effect of the virtual drop in foreign demand. Finally, the condition for a negative exchange rate elasticity of total sales for firms with trade strategy *(1,1)* can be expressed as:

$$\varepsilon^* > 1 + R \qquad (14)$$

As will be shown, the above condition (14), parameter ε, and the listed partial effects (11)–(13) can be estimated from our data on Czech manufacturing firms. So, finally, we will test the hypothesis that the terms (11)–(13) are significantly different from zero.

To proceed, we take natural logarithms from the equilibrium sales equations (7)–(10) and combine them into one equation using mutually non-exclusive dummy variables[14] $d(1,.) = d(1,0) + d(1,1)$ and $d(.,1) = d(0,1) + d(1,1)$. As a result, we get the following relationship:

$$\log[S_i(X,M)] = \log(AE') + (1\text{-}\varepsilon)\log(c) + d(1,.)\log(1+rA_xA^{-1}\tau_X^{1-\varepsilon}) + d(.,1)(1\text{-}\varepsilon)\log(r\tau_M) +$$

$$+ (\varepsilon\text{-}1)\log(e_i(M)) \qquad (15)$$

[13] As we will see below in the Results section, this assumption is consistent with our estimates.

[14] Note that using mutually exclusive trade strategy dummies would lead to the overidentification of structural parameters.

In order to convert (15) into an estimable format, let us assume that all the addends in (15) are constants[15] except the trade dummies *d(.,.)* and the productivity term $log(e_i(M))$. Furthermore, as the productivity term $log(e_i(M))$ is not directly observed, let us approximate it using an estimate of *TFP*. Given all the above, and after adding a normal i.i.d., zero-mean error term θ_{it}, equation (15) can be rewritten as follows:

$$s_{it} = \alpha_0 + \alpha_1 d(1,.)_{it} + \alpha_2 d(.,1)_{it} + \alpha_3 TFP_{it} + \theta_{it} \qquad (16)$$

where s_{it} is the log of total sales of firm *i* in time period *t*, $d(.,.)_{it}$ are dummy variables indicating trade strategies as in equation (15), and TFP_{it} is equal to $log(e_i(M))$, i.e., the firm's total factor productivity as a function of its importing strategy. The rest of the parameters of (15) are stacked into constants α_0 to α_3 of (16) as shown by the following expressions:

$$\alpha_0 = log(AE') + (1\text{-}\varepsilon)log(c)$$

$$\alpha_1 = log(1+rA_xA^{-1}\tau_X^{1-\varepsilon})$$

$$\alpha_2 = (1\text{-}\varepsilon)log(r\tau_M)$$

$$\alpha_3 = \varepsilon\text{-}1$$

which leads to:

$$\varepsilon = \alpha_3 + 1$$

$$E' = [(\alpha_3+1)/\alpha_3]^{-\alpha 3}$$

$$r\tau_M = exp(\alpha_2/\text{-}\alpha_3)$$

$$rA_x\tau_X^{1-\varepsilon} = A[exp(\alpha_1) - 1]$$

$$R = A[exp(\alpha_1)-1]/[A+ A(exp(\alpha_1)-1)] = 1 - exp(-\alpha_1)$$

Furthermore, based on (11), (12), and (13), we can express the elasticities of a hypothetical 1% change in the value of the foreign currency vis-à-vis sales on market *j*, $\rho_j(X,M)$, in terms of the estimates of (16):

$$\rho_0(0,1) = \rho(0,1) = \rho_0(1,1) = -\alpha_3 \qquad (17)$$

$$\rho_x(1,0) = 1$$

$$\rho_x(1,1) = 1 - \alpha_3 \qquad (18)$$

$$\rho(1,1) = 1 - \alpha_3 - exp(-\alpha_1) \qquad (19)$$

[15] Note that some of the assumptions about these constants could be relaxed and made firm-specific or time-variant. For example, the term $rA_xA^{-1}\tau_X^{1-\varepsilon}$, i.e., the trade-cost weighted ratio of the foreign demand level to the domestic demand level could be firm-specific based on the firm's exposure to foreign markets and the mix of foreign countries in the portfolio of the firm. Similarly, the productivity mark-up dummy for using imported intermediates, $e_i(M)$, could be continuous based on the share of imported goods in total intermediate products used. This would allow us to derive firm-specific exchange rate elasticities. This interesting extension is beyond the scope of the present paper and is left for future research.

Following our assumptions in the model, we expect α_0, α_1, and α_3 to be positive and α_2 to be negative. Regarding the estimable structural parameters of interest, we expect $\varepsilon > 1$, $r\tau_M > 1$, and $0 < R < 1$. Furthermore, based on the model's predictions for $\rho_j(X,M)$, we anticipate a negative $\rho_0(1,1)$ and a positive $\rho_x(1,1)$. Lastly, we are not able to predict the sign of $\rho(1,1)$ without making further assumptions about the model's parameters.

5. Data

Our sample consists of an unbalanced panel of 7,356 Czech manufacturing firms. The motivation to focus on the time period from 2003 to 2006 will be explained in more detail in the next paragraphs. The dataset was obtained from the Albertina database collected by the private company Creditinfo Czech Republic, s.r.o., which is available at the Czech National Bank. Although several commercial firm databases exist in the Czech Republic, to our knowledge only Albertina contains information on exports and imports.

One of the key advantages of analyzing the exports and imports of Czech firms during the said period arises from the entry of the Czech Republic into the EU in 2004. EU entry represents an exogenous event for firms and is associated with the lifting of trade barriers within the union. This implies that several non-trading Czech firms were able to participate in international trade after 2004 due to lower fixed and variable costs of accessing foreign markets. Table 1 shows the tendency of several firms shifting toward exporting and importing strategies in our sample after 2004. In particular, the share of firms that both export and import, denoted by the dummy variable $d(1,1)$, increases from about 25% in 2003 and 2004 to around 40% in 2005 and 2006. For additional firm-level and macro evidence on high trade intensity in the Czech Republic see the export and import ratios in Table A1 and Table A9 in the Appendix.

As our panel is unbalanced, we also checked whether the increased share of exporters and importers stems from trade strategy switchers or new entrants to the dataset. We are mostly interested in switchers, since our main results – the model-implied exchange rate elasticities – are functions of export and import dummy coefficient estimates.[16] This is because switchers allow us to identify these dummy coefficients from within-firm variation in trade strategies after controlling for firm-specific fixed effects. Given the time period analyzed, within-firm variation in trade strategies is likely to be associated with exogenous EU entry. It turned out that more than 14% of the observations in the pooled sample are firms that switched their trade strategy compared to the preceding year.

Further stylized facts are consistent with the hypothesis of the lifting of trade barriers implied by EU entry. According to the last column of the first row in Table 2, more than 48% of trade strategy shifts depart from a no-trade status quo. Next, according to the last row of column $d(1,1)$ in Table 2, up to 47% of trade strategy shifts lead to strategy $d(1,1)$ of both exporting and importing. At the same time, Table 3 shows that roughly 70% of the observations in the pooled sample consist of firms not switching their trade strategy of no-trade $d(0,0)$ or full trade $d(1,1)$ compared to the preceding year. This suggests that many firms cannot access foreign markets, but once a firm manages to export and import, it will tend to stay with that strategy. In other words, we observe substantial persistence in trade strategies on the micro-level, which may imply sunk fixed costs associated with those strategies.[17]

[16] See the sales equation (16).

[17] Roberts and Tybout (1997) find similar persistence patterns in the exporting activities of Colombian firms.

One of the key building blocks of the model in Section 3 was the productivity or firm size ranking by trade strategies. Firms not engaging in trade were the smallest, least productive ones, and firms both exporting and importing were the largest, most productive ones. We looked at the descriptive statistics by trade strategy sub-samples indicated by the mutually exclusive dummy variables d(export,import) to check the consistency of the data with the model. For standard descriptive statistics of variables associated with firm size, see Table A1 in Appendix 1.

To test whether there are statistically significant differences in indicators X_{it} across trade strategy sub-samples compared to the baseline case of no trade we follow Kasahara and Lapham (2013). This means estimating the trade dummy coefficients in the equation below by OLS on the pooled sample and also using fixed effects. Note that the latter estimator focuses on within-firm variation, which is our key variable of interest. Vector Z_{it} contains year dummies and, in the case of pooled OLS, also industry dummies. The term ω_{it} is assumed to be an i.i.d. normal disturbance.

$$logX_{it} = a_0 + a_1d(0,1)_{it} + a_2d(1,0)_{it} + a_3d(1,1)_{it} + A_4Z_{it} + \omega_{it}$$

The estimates of the above equation can be found in Table 4. The vast majority of the dummy coefficients are significantly different from zero, suggesting positive log-premia in the indicators for the trading strategies. Comparing the coefficients across the dummies as well as the standard descriptive statistics in Table A1 in Appendix 1, we find consistency with the model's assumptions in most cases.[18]

Table 1: Percentage of Firms in Trade Strategies d(Export,Import) by Year

	2003	2004	2005	2006
d(0,0)	58	63	42	44
d(1,0)	12	10	8	7
d(0,1)	5	4	8	10
d(1,1)	26	22	42	39
Total	100	100	100	100

Table 2: Percentage Shares of Trade Strategy Switches in 2003–2006

To strategy: From strategy:	d(0,0)	d(1,0)	d(0,1)	d(1,1)	Total switches
d(0,0)		12.1	13.8	22.2	48.1
d(1,0)	5.7		0.3	17.2	23.1
d(0,1)	4.6	0.3		7.1	12.0
d(1,1)	5.3	4.9	6.7		16.8
Total switches	15.6	17.2	20.7	46.6	100.0

Note: The total number of switches during the 2003–2006 period equals 2,630.

[18] The purpose of the exercise was merely to describe the data and to perform a consistency check of the model's assumptions. Therefore, the estimates in Table 4 should be interpreted as stylized facts without the aim to test causal relationships. In the latter case we would have had to specify other firm characteristics as explanatory variables.

Table 3: Percentage Shares of Transitions Between Trade Strategies in 2003–2006

To strategy: / From strategy:	d(0,0)	d(1,0)	d(0,1)	d(1,1)	Total
d(0,0)	38.8	2.9	3.3	5.4	50.4
d(1,0)	1.4	4.6	0.1	4.2	10.2
d(0,1)	1.1	0.1	3.3	1.7	6.2
d(1,1)	1.3	1.2	1.6	29.1	33.2
Total	42.6	8.7	8.3	40.4	100.0

Note: The total number of switches during the 2003–2006 period equals 2,630.

Table 4: Log Premia of Trade Strategies d(Export,Import) Compared to Non-Traders (2003–2006)

	Pooled OLS			Fixed effects		
Natural logarithms of indicators X	d(1,0)	d(0,1)	d(1,1)	d(1,0)	d(0,1)	d(1,1)
Sales	1.267***	1.732***	2.627***	0.063***	0.076***	0.130***
	(0.037)	(0.041)	(0.024)	(0.012)	(0.013)	(0.011)
Real value added	1.281***	1.486***	2.452***	0.067***	0.094***	0.106***
	(0.037)	(0.041)	(0.023)	(0.015)	(0.016)	(0.014)
Real capital	1.725***	1.934***	3.317***	0.035*	0.043**	0.083***
	(0.055)	(0.061)	(0.035)	(0.020)	(0.021)	(0.017)
Labor	1.187***	1.046***	2.075***	0.046***	0.018	0.058***
	(0.033)	(0.037)	(0.021)	(0.017)	(0.018)	(0.015)
Real energy and materials	1.201***	1.580***	2.634***	0.110***	0.094***	0.192***
	(0.053)	(0.058)	(0.034)	(0.023)	(0.024)	(0.020)
Real value added per labor	0.094***	0.440***	0.378***	0.021	0.076***	0.048***
	(0.021)	(0.024)	(0.014)	(0.020)	(0.021)	(0.017)
Observations		18344			18344	
Firms		7356			7356	

Note: Real values represent constant prices of 2005. Year dummies were included in all regressions. Industry dummies were included in the pooled OLS regressions. Standard errors are reported in parentheses; *, **, and *** denote significance at the 90, 95, and 99% levels.

6. Estimation

In this section we describe the estimation of equation (16), which involves three main issues. First, the variable TFP_{it}, firm i's total factor productivity as a function of its importing strategy, is fitted from a production function separately in subsection 6.1. Second, as firms select into trade strategies $d(X,M)_{it}$ endogenously, we have to correct the estimates of α_1, α_2, and α_3 for the probability of being in the respective strategies. The endogeneity of trade strategy selection follows from our model, where firms choose a trade strategy depending on their current productivity (*TFP*) and sector-specific fixed and variable costs associated with trade. Therefore, current period realizations of the sector- and firm-specific cost parameters left in the error term θ_{it} may be correlated with dummies $d(0,1)_{it}$, $d(1,0)_{it}$, and $d(1,1)_{it}$. The probabilities of choosing different trade strategies are estimated from a multinomial probit model in subsection 6.2. The third estimation issue relates to the potential correlation of TFP_{it} with the error term θ_{it}, which is the current period realization of the sales shock. This can lead to a biased estimate of α_3 if it not instrumented. The solution to this issue is briefly described in subsection 6.3.

6.1 Estimation of the Production Function

Regarding the estimation of *TFP* as a function of the importing strategy, we consider a standard Cobb-Douglas production function extended to include imported inputs as an additional factor of production:

$$y_{it} = \beta_0 + \beta_1 k_{it} + \beta_2 l_{it} + \beta_3 d(.,1)_{it} + \omega_{it} + \eta_{it} \qquad (20)$$

where y_{it} is log real value added, k_{it} is the log of the real capital stock, l_{it} is the log of the number of employees,[19] $d(.,1)_{it} = d(0,1)_{it} + d(1,1)_{it}$ is a dummy variable indicating the use of imported intermediates, ω_{it} is an unobserved firm-specific productivity shock, and η_{it} is an i.i.d. error term from the normal distribution. As the unobserved productivity shock ω_{it} is correlated with the factor inputs and the import dummy, the OLS estimates of β_0 to β_3 are, in general, biased. To solve this endogeneity issue, we combine several approaches available in the literature.

A general method of moments solution to the above endogeneity problem in the context of panel data is offered by Blundell and Bond (1998), among others. Their method, however, involves lagged dependent variables and first differencing, which may result in a weak instrument problem, erode sufficient variation, and worsen potential measurement errors in the explanatory variables, as also noted by Galuščák and Lízal (2011).

Olley and Pakes (1996), abbreviated as OP henceforth, take a different approach by approximating the productivity shock ω_{it} using investment as a proxy variable. They estimate the production function in two steps. The first step focuses on identifying the productivity shock. The second step involves instrumenting for the freely variable input, labor, via GMM and assuming capital to be predetermined. The OP method was also applied in the context of imported inputs included in the production function by Halpern et al. (2011). Levinsohn and Petrin (2003), abbreviated as LP further on, criticize the OP approach, arguing that one observes many zero investment cases in firm-level datasets, possibly due to non-convex adjustment costs. This can result in inefficient estimates and a weak proxy problem. LP approximate the productivity shock ω_{it} using energy and material inputs instead of investment and estimate the production function in two steps, similarly to OP. Kasahara and Rodrigue (2008) extend the framework of LP by adding imported intermediates to the production function as an additional predetermined variable next to capital. A further extension of the LP procedure can be found in Galuščák and Lízal (2011), who propose to correct the measurement error in real capital by means of further instruments, such as the depreciation rate, employment, and gas consumption.

Wooldridge (2009) suggests an improvement in the LP procedure allowing the estimation of the production function (20) in one step, i.e., more efficiently. The procedure requires one to assume that the error term η_{it} is uncorrelated with all of the factor inputs and their lags.[20] Furthermore, the dynamics of the unobserved productivity shock are also somewhat restricted. Galuščák and Lízal (2011) also perform measurement error correction in real capital using the Wooldridge (2009) approach and conclude that the correction yields considerably higher coefficients of real capital, just as in the LP case. In our paper, we estimate the production function following Wooldridge (2009), which is simpler than LP. We also correct for the measurement error in real capital, similarly to Galuščák and Lízal (2011). In addition, we extend the production function to include a binary indicator of imported intermediates, based on Kasahara and Rodrigue (2008), and consider firm-specific fixed effects.

[19] A more commonly used measure of the labor input, hours worked, is not available in our dataset.

[20] The same is not assumed about the unobserved productivity shock ω_{it}.

Next we outline our estimation procedure based on elements of Wooldridge (2009), Kasahara and Rodrigue (2008), and Galuščák and Lízal (2011). Suppose that material inputs m_{it} depend on capital, the import dummy, and the productivity shock ω_{it}:

$$m_{it} = f(k_{it}, \omega_{it}, d(.,1)_{it}) \quad (21)$$

and f is an invertible and monotonic function of ω_{it}, so that we can write:

$$\omega_{it} = g(k_{it}, m_{it}, d(.,1)_{it}) \quad (22)$$

Assume that the error term η_{it} is uncorrelated with the current values and lags of labor, capital, the import dummy, and material inputs m_{it}:

$$E(\eta_{it} \mid l_{it}, k_{it}, d(.,1)_{it}, m_{it}, \ldots, l_{i1}, k_{i1}, d(.,1)_{i1}, m_{i1}) = 0 \quad (23)$$

The dynamics of the unobserved productivity shock are restricted as:

$$E(\omega_{it} \mid k_{it}, d(.,1)_{it}, l_{it-1}, k_{it-1}, d(.,1)_{it-1}, m_{it-1}, \ldots) = E(\omega_{it} \mid \omega_{it-1}) =$$

$$= j(\omega_{it-1}) = j(g(k_{it-1}, m_{it-1}, d(.,1)_{it-1}) \quad (24)$$

For productivity innovations a_{it} we can write:

$$\omega_{it} = j(\omega_{it-1}) + a_{it} \quad (25)$$

where

$$E(a_{it} \mid k_{it}, d(.,1)_{it}, l_{it-1}, k_{it-1}, d(.,1)_{it-1}, m_{it-1}, \ldots) = 0 \quad (26)$$

which implies that the freely variable labor and material inputs l_{it} and m_{it} are correlated with productivity innovations a_{it}, but capital k_{it}, the import dummy $d(.,1)_{it}$, and all lags of l_{it}, m_{it}, k_{it}, and $d(.,1)_{it}$ are uncorrelated with a_{it}. After plugging (24) and (25) into the production function (20) we get:

$$y_{it} = \beta_0 + \beta_1 k_{it} + \beta_2 l_{it} + \beta_3 d(.,1)_{it} + j(g(k_{it-1}, m_{it-1}, d(.,1)_{it-1})) + u_{it} \quad (27)$$

where $u_{it} = a_{it} + \eta_{it}$ and

$$E(u_{it} \mid k_{it}, d(.,1)_{it}, l_{it-1}, k_{it-1}, d(.,1)_{it-1}, m_{it-1}, \ldots) = 0 \quad (28)$$

Before estimating (27) we need to specify functions j and g. Copying the approaches used in the literature, we assume the productivity process j to follow a random walk with drift, so that (25) can be rewritten as:

$$\omega_{it} = \psi + \omega_{it-1} + a_{it} \quad (29)$$

Regarding function g, we use a third-order polynomial approximation suggested by Petrin, Poi, and Levinsohn (2004) and Wooldridge (2009):

$$\omega_{it} = g(k_{it}, m_{it}, d(.,1)_{it}) =$$

$$= h(k_{it}, m_{it}, d(.,1)_{it}, k_{it}m_{it}, k^2_{it}, m^2_{it}, k^2_{it}m_{it}, k_{it}m^2_{it}, k^3_{it}, m^3_{it}) \quad (30)$$

where h is a linear function. Using (29) and (30) we can rewrite (27) as:

$$y_{it} = (\beta_0 + \psi) + \beta_1 k_{it} + \beta_2 l_{it} + \beta_3 d(.,1)_{it} + g(k_{it-1}, m_{it-1}, d(.,1)_{it-1}) + u_{it} \quad (31)$$

Note that in (31) we end up including a learning-by-importing effect via the lagged import dummy $d(.,1)_{it-1}$ as in Kasahara and Rodrigue (2008).

Next, we estimate (31) by GMM and two-stage least squares, treating labor l_{it} as endogenous, correcting for the measurement error in capital k_{it} and assuming $d(.,1)_{it}$ to be predetermined given the approximation for ω_{it}. In both estimation methods we use lagged labor l_{it-1}, the log of depreciation, and the log of energy and material inputs m_{it} as instruments, similarly to Wooldridge (2009) and Galuščák and Lízal (2011). In the two-stage least squares version we also assume firm-specific fixed effects, which turn out to be important.

After fitting the production function (31), we save the estimate of total factor productivity in natural logarithm (*tfp*) as a function of the import strategy. This means recording the following expression:

$$tfp_{it} = y_{it} - \beta_1 k_{it} - \beta_2 l_{it} \quad (32)$$

This expression is used in the remaining stages of our estimation, i.e., the multinomial probit models of trade strategy choice and the equilibrium sales equation.

6.2 Estimation of the Probabilities of Choosing Trade Strategies

To address the problem of non-random samples of firms selecting into different trade strategies in equation (16), we estimate the probabilities of choosing each of the four trade strategies using a year-by-year multinomial probit model. The firm- and year-specific probabilities will be then used as instruments for dummy variables $d(1,.)_{it}$, $d(.,1)_{it}$ in equation (16). The multinomial probit approach is motivated by the unobserved ordering of trade strategies. As noted in section 3, trade strategy choice is determined by firm i's productivity parameter e_i and the cut-off productivities for each strategy depending on the relative slopes of trade strategy-specific equilibrium profit functions $\Pi_i*(X,M)$. Using the multinomial probit we do not have to make further assumptions about the parameters of $\Pi_i*(X,M)$.

Trade strategy choice in the multinomial probit framework is modeled as follows. We introduce latent variables γ_{ij} indexed for each firm i and trade strategy choices j from the set $(X,M) = \{(0,0), (0,1), (1,0), (1,1)\}$ and consider a 1 x q row vector of exogenous firm-specific variables $\boldsymbol{w_i}$:

$$\gamma_{ij} = \boldsymbol{w_i}\, \delta_j + \xi_{ij}$$

where ξ_{iX}, ξ_{iM}, and ξ_{iXM} are distributed independently and identically standard normal. The firm chooses trade strategy k such that $\gamma_{ik} \geq \gamma_{im}$ for $m \neq k$. Taking the difference between γ_{ik} and γ_{im} we get:

$$\Gamma_{i,k,m} = \gamma_{ik} - \gamma_{im} = \boldsymbol{w_i}\, (\delta_k - \delta_m) + (\xi_{ik} - \xi_{im}) = \boldsymbol{w_i}\varphi_{k'} + \omega_{ik'}$$

where $Var(\omega_{ik'}) = Var(\xi_{ik} - \xi_{im}) = 2$ and $Cov(\omega_{ik'},\omega_{il'}) = 1$ for $k' \neq l$. Using the above expressions we can write the probabilities of choosing each of the four trade strategies as follows:

$$Prob(i\ chooses\ (0,0)) = Prob(\Gamma_{i,00,01} \geq 0, \Gamma_{i,00,10} \geq 0, \Gamma_{i,00,11} \geq 0)$$

$$Prob(i\ chooses\ (1,0)) = Prob\ (\Gamma_{i,10,00} \geq 0, \Gamma_{i,10,01} \geq 0, \Gamma_{i,10,11} \geq 0)$$

$$Prob(i\ chooses\ (0,1)) = Prob(\Gamma_{i,01,00} \geq 0, \Gamma_{i,01,10} \geq 0, \Gamma_{i,01,11} \geq 0)$$

$$Prob(i\ chooses\ (1,1)) = Prob(\Gamma_{i,11,00} \geq 0, \Gamma_{i,11,01} \geq 0, \Gamma_{i,11,10} \geq 0)$$

The above probabilities indicate that choice in the multinomial probit model is based on the multivariate normal distribution $MVN(0,\boldsymbol{\Sigma})$, where $\boldsymbol{\Sigma}$ is a 3 x 3 variance-covariance matrix with 2-s on the diagonal and 1-s off the diagonal.

We estimate the year-by-year multinomial probits as defined above with exogenous firm-specific variables $\boldsymbol{w_i}$ including the log of capital approximating firm size, *tfp* as a function of importing from (32), a dummy for foreign ownership, a lagged trading dummy indicating engagement in any of the trade strategies except (0,0) in the preceding period,[21] and a set of industry dummies. As a concluding step, the fitted probabilities for each firm and time period are recorded.

6.3 Estimation of the Equilibrium Sales Equation

Once tfp_{it} in (32) and the trade strategy probabilities have been fitted, all that remains is to estimate the equilibrium sales equation (16). We apply two-stage least squares to instrument for the export and import dummies $d(1,.)_{it}$ and $d(.,1)_{it}$ using the firm- and year-specific fitted probabilities associated with the dummies as instruments. We also consider firm-specific fixed effects in sales. Finally, we perform linear and non-linear tests of combinations of the sales equation's coefficient estimates. This allows us to test some of the model's structural parameters and the implied exchange rate elasticities in (17)–(19), as presented in Table 9 in the next section.

7. Results

Table 5 presents estimates of the production function based on several approaches. Columns (1)–(4) follow and extend the frameworks of Wooldridge (2009) and Galuščák and Lízal (2011) and deal with endogenous variables via GMM. Column (1) is the replication of Wooldridge (2009) on our Czech sample. This involves estimating equation (31) by GMM and treating labor as endogenous. The estimates in column (2) result from the extension of Wooldridge (2009) as suggested by Galuščák and Lízal (2011). The latter authors suggested a measurement error correction in capital using, for example, depreciation and energy inputs as instruments apart from the treatment of endogenous labor. The models in columns (3) and (4) extend the specifications used in (1) and (2) to include an import dummy, which is assumed to be exogenous given the proxy for the productivity shock in the same period, as suggested in equation (30).

Comparing our estimates in columns (1) and (2) with those of Galuščák and Lízal (2011) we find similar results. Specifically, correcting the measurement error in capital is important, as the log capital coefficient

[21] The indicator of prior trade experience is important given the observed persistence in trade strategies in our dataset. Past exporting activities were found to be a good predictor of future engagement in exports also by Roberts and Tybout (1997) based on a sample of Colombian firms.

increases sharply after the correction. At the same time, the elasticity of labor remains roughly the same. However, the sizes of the estimated coefficients are different in the two studies. This may be largely due to the fact that we use the number of employees instead of hours worked as the proxy for labor. Our choice of the number of employees was predetermined by data limitations.

The last four columns of Table 5 present results from the models including firm-specific fixed effects, and endogenous variables are treated by two-stage least squares. The specifications and the pattern of treating endogenous variables are the same as in the first half of Table 5. Specifically, in the column (5) model, labor is instrumented but the measurement error in capital is not corrected. In the column (6) estimates, the measurement error in capital has been instrumented by depreciation and energy and material costs. Columns (7) and (8) replicate the latter two columns while also including the import dummy.

Comparing the results in the two halves of the table, all the coefficient estimates are roughly halved but remain statistically significant after considering firm-specific fixed effects. This implies that fixed effects are likely to be endogenous and therefore should not be disregarded in similar studies.

Regarding the coefficient on the import dummy, the estimate of key interest to us within the production function, we can say that imported intermediates tend to increase total factor productivity significantly. However, after the measurement error in capital has been corrected, the effect of imported intermediates is roughly halved. The same conclusion holds for both the GMM and the 2SLS fixed effects estimates. To sum up, the above results are in line with the assumptions made in our model and similar to other studies considering import dummies in the production function, such as Kasahara and Rodrigue (2008).

As we have concluded that both firm-specific fixed effects and the measurement error correction in capital are important, we will use the estimated TFP based on column (8) in what follows. Note that during the production function estimation we were forced to work with a reduced sample due to data limitations. This meant considering only 4,815–5,180 different firms instead of the full sample of 7,356 firms depending on the method of estimation and the associated data requirements. However, to recover a TFP estimate for each firm, we only need to observe labor and capital and use the associated coefficient estimates. Thanks to this fact we can also estimate TFP out of the production function sample. Therefore, as a sensitivity check we will replicate the final results of our study for both the *full* and the *reduced* sample. By *full sample* we mean the sample also containing TFP estimates out of the sample considered for estimating the production function. Similarly, when referring to the *reduced sample* we mean keeping only those observations which were used in the production function estimation.

The fitted TFP from above first enters the estimation of the probabilities of being in a particular trade strategy from the year-by-year multinomial probit models. To keep the summary of results to a manageable size, we present estimates only for the pooled sample in 2003–2006 in Table 6. For the year-by-year estimates we refer the interested reader to Tables A2–A5 in Appendix 2.

The coefficients on log real capital and log TFP in Table 6 suggest that an increase in these variables improves the probabilities of being in any of the trading strategies compared to the base outcome of no trade. The coefficients of these two regressors tend to be the largest for the full trade strategy d(1,1), which implies that any increase in the two regressors increases the probability of being in full trade the most. The findings thus do not contradict our model in general. Furthermore, foreign ownership tends to increase the probability of a firm being involved in international trade. The size of the coefficient on the foreign ownership dummy, however, does not follow a clear systematic pattern over time and across different trade strategies. The coefficient on the lagged trade dummy is significantly positive, which suggests persistence in trade strategies.[22] We can also assert this because once a firm starts trading, it is likely to stick to this strategy afterwards. Finally, we can observe some systematic patterns in the coefficients on the listed industry dummies, though interpreting them is not the main focus of the present study.

After obtaining the fitted firm- and year-specific *TFP* and the probabilities of being in a particular trade strategy, we estimated the sales equation. This allows us to identify selected structural parameters of the model and to estimate the exchange rate elasticities of sales. The estimates of the sales equation itself, for both the full and the reduced samples, can be found in Table 7 below. The signs of the export and import dummy coefficients and log *TFP* are as expected and in accordance with our model in both samples. Unfortunately, though, the coefficient estimate of the import dummy is insignificant in both versions of the dataset.[23] Note, however, that the imprecise estimate of α_2 in (16) only affects the estimate of the structural parameter $r\tau_M$ (Table 8) discussed below and does not influence our main results regarding the exchange rate elasticities (Table 9).

By using the estimates of the sales equation in Table 7 we can derive estimates of some of the model's structural parameters. These are summarized and tested in Table 8. The estimate of the elasticity of substitution ε is greater than one and thus is in accordance with the theory. The estimated share of the freight cost-discounted foreign demand level in the total demand level faced by exporter firms, *R*, lies between zero and one as expected. The product of the unit cost of importing and the nominal exchange rate $r\tau_M$ exceeds one, which is again in line with the model's assumptions. Notably, there are some differences between the three estimates depending on whether the full or the reduced sample is used, especially in the case of parameter $r\tau_M$. Moreover, the standard error of the latter estimate is relatively large, making the point estimate indistinguishable from zero. This is likely to be a result of the imprecise estimate of coefficient α_2 in the sales equation (16).

In addition to the above structural parameters of the model we can use the estimates of the sales equation (16) to express the exchange rate elasticities of sales as predicted by the model. The elasticities tell us the percentage response of sales to the nominal exchange rate depreciating by one percent. As the elasticities are symmetric with respect to a positive or a negative currency shock, we present the elasticities of the opposite sign to look at the response of sales to the appreciation of the domestic currency in Table 9 below. This is motivated by the fact that appreciation shocks usually get more attention in Czech economic news reports.

According to our results in Table 9, a one percent appreciation of the domestic currency leads to a 0.2% rise in domestic sales for firms which import some of their inputs. The same shock causes export sales to drop by 1% if the firm does not import inputs, as the exporters are assumed to be price takers on foreign markets and export sales are assumed to be contracted in foreign currency. The similarly negative impact on export sales is somewhat reduced to 0.8% if the firm uses imported intermediate goods. In the case of total sales of firms that both export and import, the appreciation shock leads to a drop of 0.2% or 0.4%, depending on whether the estimate is based on the full or the reduced sample. The above elasticity estimates are roughly comparable to our estimates on macro data. For more details on the macro estimates see Appendix 3.

[22] Persistence in trading activities is consistent with the findings of Roberts and Tybout (1997) on Colombian firm-level data.

[23] The reason for the above result is probably the fact that the two trade dummies in equation (16) are correlated.

Table 5: Estimates of the Production Function

The dependent variable is the log of real value added. Estimation period: 2003–2006.

Estimator:	GMM				IV-2SLS with fixed effects			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Constant	5.644***	3.867***	5.266***	3.655***	7.899***	2.876***	7.858***	2.895***
	(0,474)	(0,871)	(0,48)	(0,858)	(0,43)	(0,578)	(0,429)	(0,576)
Log of the number of employees	0.458***	0.426***	0.452***	0.422***	0.213***	0.287***	0.216***	0.287***
	(0,015)	(0,019)	(0,015)	(0,019)	(0,039)	(0,05)	(0,039)	(0,05)
Log of real capital	0.261***	1.528***	0.254***	1.489***	0.185***	0.760***	0.183***	0.756***
	(0,021)	(0,141)	(0,021)	(0,138)	(0,011)	(0,034)	(0,011)	(0,034)
Import dummy d(0,1) + d(1,1)	-	-	0.205***	0.099***	-	-	0.073***	0.039**
			(0,017)	(0,024)			(0,013)	(0,017)
R-squared	0,829	0,635	0,832	0,648	0,809	0,760	0,813	0,762
Number of observations	12434	11393	12434	11393	12434	11393	12434	11393
Number of firms	5180	4815	5180	4815	5180	4815	5180	4815

Notes: Standard errors are in parentheses. *, **, and *** denote significance at the 90, 95, and 99% levels. Year dummies were included in all regressions.

Estimates:
(1) follows Wooldridge (2009)
(2) Wooldridge (2009), real capital instrumented by depreciation and energy and material inputs
(3) Wooldridge (2009), import dummy included
(4) Wooldridge (2009), import dummy included and real capital instrumented by depreciation and energy and material costs
(5) IV-2SLS version of Wooldridge (2009) also including fixed effects
(6) IV-2SLS version of Wooldridge (2009) also including fixed effects; capital instrumented by depreciation and energy and material costs
(7) IV-2SLS version of Wooldridge (2009) also including fixed effects and the import dummy
(8) IV-2SLS version of Wooldridge (2009) also including fixed effects and the import dummy; capital instrumented by depreciation and energy and material costs

Table 6: Estimates of the Multinomial Probit Model of Trade Strategy Choice

Estimates by choice outcomes d(export,import) and d(0,0) as the base outcome.
Estimation interval: 2003–2006.

Choice outcomes:	d(1,0)	d(0,1)	d(1,1)
Constant	-3.782***	-5.065***	-7.069***
	(0.227)	(0.323)	(0.257)
Log real capital	0.211***	0.212***	0.458***
	(0.009)	(0.009)	(0.009)
Log TFP	0.147***	0.206***	0.228***
	(0.012)	(0.013)	(0.011)
Foreign ownership dummy	0.657***	0.502***	0.497***
	(0.130)	(0.141)	(0.125)
Lagged trade dummy	1.640***	1.487***	2.176***
	(0.042)	(0.044)	(0.037)
Light industry dummy	-0.678***	0.428	0.354
	(0.206)	(0.308)	(0.238)
Raw materials industry dummy	-0.405**	0.482	0.444*
	(0.206)	(0.308)	(0.238)
Machinery industry dummy	-0.042	0.458	0.813***
	(0.209)	(0.311)	(0.240)
Electric industry dummy	-0.730***	0.534*	0.697***
	(0.212)	(0.311)	(0.241)
Car manufacturing industry dummy	-0.614***	0.290	0.900***
	(0.232)	(0.328)	(0.252)
Number of observations		20165	

Notes: Standard errors are in parentheses. *, **, and *** denote significance at the 90, 95, and 99% levels. The above model was estimated on the pooled sample of 2003–2006 with the largest number of observations. In further estimation we use fitted choice probabilities estimated from year-by-year multinomial probit models. The year-by-year estimates of the model can be found in Appendix 2.

Table 7: Estimates of the Equilibrium Sales Equation

The dependent variable is the log of total sales.

	Coefficients of eq. (16)	Full sample	Reduced sample
Constant	α_0	3.666***	3.989***
		(0.000)	(0.000)
Export dummy d(1,0)+d(1,1)	α_1	0.585**	0.907**
		(0.000)	(0.000)
Import dummy d(1,0)+d(1,1)	α_2	-0.008	-0.208
		(0.000)	(0.000)
Log TFP as a function of import dummy	α_3	0.201***	0.227***
		(0.000)	(0.000)
R-squared		0.077	0.053
Number of observations		18344	11217
Number of firms		7356	4752

Note: The equation was estimated by 2SLS including fixed effects. Log TFP was fitted from the production function in Table 5, column 8. Standard errors are reported in parentheses. *, **, and *** denote significance at the 90, 95, and 99% levels. The reduced sample corresponds to the observations used in Table 5, column 8.

Table 8: Estimates of Selected Structural Parameters

	Parameter in the model	Coefficients of eq. (16)	Full sample	Reduced sample
Elasticity of substitution of the CES utility function	ε	$1 + \alpha_3$	1.201***	1.227***
			(0.072)	(0.073)
Share of the freight cost-discounted foreign demand level in the total demand level faced by	R	$1 - \exp(-\alpha_1)$	0.443***	0.597***
			(0.148)	(0.159)
Variable unit cost of imports (CZK thousands)	$r\tau_M$	$\exp(\alpha_2/-\alpha_3)$	1.042	2.501
			(0.929)	(2.505)
Number of observations			18344	11217
Number of firms			7356	4752

Note: Standard errors are reported in parentheses and are obtained by the delta method in the case of the last two parameters. *, **, and *** denote significance at the 90, 95, and 99% levels.

Table 9: Implied Exchange Rate Elasticities of Sales

(% change in sales / domestic currency appreciating by 1 %)	Model	Coefficients of eq. (16)	Full sample	Reduced sample
Domestic sales in strategies d(1,1) and d(0,1)	$-\rho0(0,1) = -\rho(0,1) =$	α_3	0.201***	0.227***
	$= -\rho0(1,1)$		(0.072)	(0.073)
Export sales in strategy d(1,1)	$-\rho_x(1,1)$	$\alpha_3 - 1$	-0.799***	-0.775***
			(0.072)	(0.075)
Total sales in strategy d(1,1)	$-\rho(1,1)$	$\alpha_3 + \exp(-\alpha_1) - 1$	-0.243*	-0.370**
			(0.127)	(0.161)
Number of observations			18344	11217
Number of firms			7356	4752

Note: Standard errors are reported in parentheses. The delta method is used to obtain the standard error in the case of the last elasticity. *, **, and *** denote significance at the 90, 95, and 99% levels.

8. Conclusion

We studied the impact of a hypothetical currency shock on firm sales depending on a mix of firms' exporting and importing strategies. We argue that the exchange rate pass-through to sales is special in the case of firms that both export and import, a class of firms that became more widespread after the Czech Republic entered the European Union. Accordingly, we used within-firm variation in the time period around EU entry to identify our estimates. Our aim was to capture the exogenous effect of the lifting of trade barriers associated with EU entry on the participation of firms in international trade.

We found that importing firms are partially able to cushion the negative impact of an exchange rate shock on their export sales. In particular, the drop in export sales as a result of the domestic currency appreciating by one percent is 0.8% if the firm imports some of its intermediate goods, instead of 1% if a price taker firm does not import inputs. The above elasticities of export and total sales are roughly in line with our estimates on macro-level data.

The topic of exchange rate shocks has received heightened attention in the policy sphere and media recently, as the Czech National Bank decided to weaken the Czech koruna by starting to carry out interventions on the foreign exchange markets for an undefined period in November 2013. Regarding this policy shock to the Czech koruna, our findings suggest that the benefits from the recently improved price competitiveness of Czech exporters will be somewhat dampened if exporters have to rely on more expensive imported inputs.

Our analysis contributes to the literature on heterogeneous firms and trade by studying the impact of a hypothetical exchange rate shock on firm sales, an issue which has not been studied before in this context. Furthermore, our paper offers a simple static alternative to the dynamic model of exporting and importing with heterogeneous firms by Kasahara and Lapham (2013). In contrast to the above paper we get testable implications that are easy to estimate. Next, as opposed to Bas and Strauss-Kahn (2011), we test the model's implications by estimating the equilibrium sales equation obtained directly from the model. Finally, we confirm that imported intermediates increase the total factor productivity of firms, as found also by Bas and Strauss-Kahn (2011), Halpern et al. (2011), and Kasahara and Rodrigue (2008) on micro data from France, Hungary, and Chile, respectively.

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9. Appendix 1: Descriptive Statistics

Table A1: Descriptive Statistics by Trade Strategies d(Export,Import) in 2003–2006

		d(0,0) no exports no imports	d(1,0) exports no imports	d(0,1) no exports imports	d(1,1) exports imports	Full sample
Sales	mean	88270	167333	293281	1009744	414152
	st. dev.	(444521)	(358409)	(683406)	(5159990)	(3018112)
Real value added		19806	38958	64990	202585	85083
		(80793)	(52316)	(172312)	(814707)	(481300)
Real capital		24851	50160	105218	330732	133819
		(141595)	(119029)	(649103)	(1591098)	(946767)
Labor		57	129	122	345	163
		(115)	(163)	(231)	(865)	(526)
Energy and materials		56576	76772	142500	485441	243268
		(353289)	(223157)	(427539)	(2783210)	(1822603)
Exports		0	52296	0	468160	159252
			(151423)		(934535)	(580949)
Imports		0	0	60102	299239	103036
				(195661)	(745504)	(453120)
Real value added per labor		419	475	691	656	522
		(862)	(1325)	(1641)	(3095)	(1976)
Exports to imports		0	0	0	4,3	3,5
					(9.2)	(8.5)
Exports to sales		0	0,24	0	0,30	0,29
			(0.19)		(0.26)	(0.25)
Imports to sales		0	0	0,17	0,18	0,18
				(0.21)	(0.21)	(0.21)
Imports per energy and materials		0	0	3,1	1,1	1,5
				(20.0)	(9.6)	(12.0)
Observations		9319	1665	1306	6054	18344
Firms		4961	1130	921	2727	7356

Note: Values in thousands of Czech korunas; real values represent constant prices of 2005. Exports and imports are measured in our dataset as interval variables with values falling into one of nine categories.

10. Appendix 2: Year-by-Year Estimates of the Multinomial Probit

Tables A2-A5: Estimates of the Multinomial Probit Model of Trade Strategy Choice

The estimates by choice outcomes d(export,import) consider no trade d(0,0) as the

Sample: 2003

Choice outcomes:	d(1,0)	d(0,1)	d(1,1)
Constant	-4.175***	-6.109***	-8.467***
	(0.327)	(0.576)	(0.390)
Log real capital	0.278***	0.337***	0.640***
	(0.014)	(0.018)	(0.016)
Log TFP	0.122***	0.093***	0.094***
	(0.014)	(0.017)	(0.014)
Foreign ownership dummy	0.488***	0.342	0.379**
	(0.180)	(0.222)	(0.180)
Lagged trade dummy	1.735***	1.197***	1.633***
	(0.059)	(0.071)	(0.056)
Light industry dummy	-1.031***	0.324	0.038
	(0.295)	(0.546)	(0.354)
Raw materials industry dummy	-0.752**	0.424	0.266
	(0.296)	(0.547)	(0.355)
Machinery industry dummy	-0.234	0.416	0.588*
	(0.298)	(0.552)	(0.358)
Electric industry dummy	-0.935***	0.573	0.687*
	(0.304)	(0.551)	(0.359)
Car manufacturing industry dummy	-0.633**	0.515	0.795**
	(0.322)	(0.567)	(0.371)
Number of observations	9236	9236	9236

The estimates by choice outcomes d(export,import) consider no trade d(0,0) as the

Sample: 2004

Choice outcomes:	d(1,0)	d(0,1)	d(1,1)
Constant	-4.991***	-6.784***	-9.797***
	(0.518)	(0.713)	(0.604)
Log real capital	0.286***	0.346***	0.671***
	(0.023)	(0.026)	(0.025)
Log TFP	0.345***	0.391***	0.404***
	(0.046)	(0.050)	(0.046)
Foreign ownership dummy	0.242	0.476	0.279
	(0.280)	(0.293)	(0.275)
Lagged trade dummy	2.441***	1.859***	2.590***
	(0.088)	(0.098)***	(0.086)
Light industry dummy	-1.460***	-0.284	-0.522
	(0.438)	(0.636)	(0.521)
Raw materials industry dummy	-1.374***	-0.576	-0.715
	(0.441)	(0.639)	(0.524)
Machinery industry dummy	-0.981**	-0.648	-0.404
	(0.444)	(0.648)	(0.528)
Electric industry dummy	-1.450***	-0.273	0.046
	(0.453)	(0.646)	(0.529)
Car manufacturing industry dummy	-1.210**	-0.273	0.170
	(0.485)	(0.680)	(0.552)
Number of observations	5342	5342	5342

The estimates by choice outcomes d(export,import) consider no trade d(0,0) as the

Sample: 2005

Choice outcomes:	d(1,0)	d(0,1)	d(1,1)
Constant	-4.037***	-6.489***	-7.812***
	(0.463)	(0.649)	(0.500)
Log real capital	0.242***	0.288***	0.467***
	(0.018)	(0.017)	(0.017)
Log TFP	0.275***	0.481***	0.547***
	(0.043)	(0.046)	(0.042)
Foreign ownership dummy	0.948***	0.735***	0.543**
	(0.260)	(0.274)	(0.257)
Lagged trade dummy	0.975***	0.979***	2.406***
	(0.115)	(0.111)	(0.094)
Light industry dummy	-0.750*	0.503	0.228
	(0.390)	(0.593)	(0.436)
Raw materials industry dummy	-0.372	0.768	0.516
	(0.389)	(0.593)	(0.436)
Machinery industry dummy	-0.166	0.459	0.676
	(0.396)	(0.600)	(0.442)
Electric industry dummy	-0.924**	0.584	0.368
	(0.403)	(0.598)	(0.442)
Car manufacturing industry dummy	-1.339***	0.420	0.584
	(0.505)	(0.634)	(0.472)
Number of observations	5847	5847	5847

The estimates by choice outcomes d(export,import) consider no trade d(0,0) as the

Sample: 2006

Choice outcomes:	d(1,0)	d(0,1)	d(1,1)
Constant	-2.778***	-5.130***	-8.617***
	(0.590)	(0.649)	(0.638)
Log real capital	0.016	0.090***	0.319***
	(0.022)	(0.019)	(0.020)
Log TFP	-0.004	0.244***	0.368***
	(0.046)	(0.044)	(0.044)
Foreign ownership dummy	0.903**	0.990***	1.130***
	(0.359)	(0.343)	(0.336)
Lagged trade dummy	2.854***	2.527***	3.557***
	(0.108)	(0.094)	(0.101)
Light industry dummy	-0.074	0.868	1.391**
	(0.504)	(0.574)	(0.552)
Raw materials industry dummy	0.085	0.911	1.419***
	(0.502)	(0.573)	(0.550)
Machinery industry dummy	0.214	0.825	1.711***
	(0.511)	(0.581)	(0.557)
Electric industry dummy	-0.368	0.816	1.560***
	(0.516)	(0.582)	(0.559)
Car manufacturing industry dummy	-0.735	-0.131	1.507***
	(0.599)	(0.645)	(0.587)
Number of observations	5082	5082	5082

Notes: Standard errors are in parentheses. *, **, and *** denote significance at the 90, 95, and 99% levels. Choice outcomes are dummy variables according to trade status d(export,import).

11. Appendix 3: Exchange Rate Elasticity Estimates on Macro Data

In order to put the firm-level exchange rate elasticities into a broader context, it is interesting to compare them with their macro-level counterparts. As none are available on Czech data in the literature, we fill this gap in Appendix 3. In what follows, elasticities from a direct time-series approach and implied elasticities from a macro-level version of the structural equation (16) are estimated on macro data. The sensitivity of the results is checked for different time periods and with respect to the use of manufacturing or aggregated national accounts data. We conclude that the firm-level elasticity estimates are relatively close to those obtained on the macro level. However, the results are in general sensitive to the estimation period and the data source chosen. In addition, one should keep in mind the limited comparability between the micro and the macro estimates. Below we describe the data, the estimation approaches, the results, and the comparability of the micro and macro estimates in detail.

For estimation on the macro level we need to collect indicators of aggregate exports and output, total factor productivity, and the real exchange rate. First, quarterly exports and output data from the national accounts and for the manufacturing subsector are obtained from the Czech Statistical Office. These variables are published in constant prices and seasonally adjusted. Second, total factor productivity (*TFP*) is taken from the European Commission (EC). The EC's estimate is based on the standard production function approach and is published annually. We use the annual growth rate of TFP interpolated to quarterly frequency using a quadratic polynomial. Third, the real effective exchange rate (REER) index of the Czech koruna is retrieved from the Czech National Bank's database, where the nominal rates were deflated by relative PPIs and weighted by trade volumes in SITC categories 5–8. Here, an increase in the REER means appreciation of the domestic currency. In order to achieve stationarity and comparability with the annual micro data, year-on-year growth rates are used for all variables entering the estimation procedures below. Descriptive statistics of the macro dataset can be found in Table A9.

The exchange rate elasticities of exports and output on the macro-level are estimated by two simple approaches. The first, in (33) and (34), is an AR-X model of exports, X_t, and output, Y_t, respectively, where the real exchange rate, $REER_t$, is assumed to be an exogenous factor. The coefficients of the real exchange rate, b_1 and c_1, are considered for the direct exchange rate elasticities of exports, W_x, and output, W, as declared in (36) and (37) below. As an increase in the *REER* index means appreciation of the Czech koruna, W_x and W denote the elasticity of a 1% appreciation of the domestic currency.

The second approach, in (35), adapts equation (16) to the macro data. In particular, the export and import dummies in (16) are replaced with the ratios of exports to output and imports to output, XY_t and MY_t, respectively. The implied elasticities of exports, w_x, and output, w, are computed as in expressions (18) and (19) on the firm level. Specifically, using the coefficients of (35), we can express w_x and w as in (38) and (39) below. Similarly to W_x and W, and following Table 7 in the Results section, w_x and w denote the elasticity of a 1% appreciation of the domestic currency.

$$B_0(L)X_t = b_{00} + b_1 REER_{t-1} + \delta_t \quad (33)$$

$$C_0(L)Y_t = c_{00} + c_1 REER_{t-1} + \varsigma_t \quad (34)$$

$$A_0(L)Y_t = a_{00} + a_1 XY_t + a_2 MY_t + a_3 TFP_{t-1} + \chi_t \quad (35)$$

where lag polynomials $B_0(L)$, $C_0(L)$, and $A_0(L)$ assume the common form:

$$\Phi_0(L) = 1 - \sum_{i=1}^{q} \phi_{0i} L^i$$

Direct elasticity of exports: $W_x = b_1$ (36)

Direct elasticity of output: $W = c_1$ (37)

Implied elasticity of exports: $w_x = a_3 - 1$ (38)

Implied elasticity of output: $w = a_3 + exp(-a_1) - 1$ (39)

Equations (33)–(35) are estimated by ordinary least squares. Up to two lags in A_0, B_0, and C_0 are added in order to eliminate serial correlation in the error terms δ_t, ς_t, and χ_t, which are assumed to be zero-mean normal i.i.d. *REER* and *TFP* enter the equations in their first lags to avoid contemporaneous correlations with the errors.

The estimates from the three equations (33)–(35) are summarized in Tables A6, A7, and A8. As a sensitivity check, the tables compare the results from the national accounts and manufacturing data as well as across different time periods of estimation. The national accounts data are preferred to manufacturing in the case of the implied elasticity estimates in Table A8. This is because the explanatory variables Exports to GDP and Imports to GDP are available only from the national accounts, not for manufacturing. The above data limitation originates from the fact that manufacturing exports and output data are published in the form of a base index, while manufacturing imports are not available at all. Regarding the choice of estimation periods, the sub-interval between 2001 and 2008 is preferred to the full samples due to a better match with the estimation interval in the micro part.[23]

Tables A6 and A7 contain the results of the AR-X models (33) and (34) for the narrowed interval 2001–2008 and the full samples of observation, respectively. Noticeably, the estimates of W_x and W are somewhat sensitive to the time periods chosen. Furthermore, the differences between the estimates across the two data sources are more marked.

Table A8 lists the implied exchange rate elasticity estimates based on (35). As mentioned above, we consider the first column with the narrowed subsample and aggregated data to be the most relevant for the micro-macro comparison, while the remaining columns are presented as a sensitivity check. The implied elasticity estimates in the lower part of the first column are close to those obtained from the AR-X models

[23] The panel of firms covers the period 2003–2006.

(33) and (34). However, we cannot draw the same conclusion for the other combinations of data sources and estimation intervals presented in the remaining columns of Table A8.

To sum up, the macro-level estimates of the exchange rate elasticities of exports and output are relatively close to those obtained on the firm level, especially in the case of exports. Specifically, a 1% appreciation of the domestic currency is associated with a statistically significant drop in export dynamics of about 0.8 percentage points according to the macro data and of roughly the same value based on the micro data. Furthermore, the impact of an identical shock on aggregate output ranges from drops that are statistically not distinguishable from zero to a statistically significant rise of 0.1 percentage points. Contrary to the macro results, the micro estimates suggest a statistically significant drop in total sales of 0.4 to 0.2 percentage points.

At the same time it must be noted that the micro- and macro-level estimates of the exchange rate elasticities are not fully comparable. First, as demonstrated above, the macro estimates are relatively sensitive to the choices of estimation periods and data sources. Second, such comparison is possible only under the representative firm assumption. However, our micro estimates are associated with firms that both export and import. At the same time, a large proportion of firms represented in the aggregate data do not participate in international trade. Accordingly, a significant share of non-trading firms would help explain why the macro estimates of the exchange rate elasticity of total output are closer to zero in contrast to a significantly negative micro estimate.

Table A6: Exchange Rate Elasticity Estimates on Macro Data (AR-X model)

All variables in year-on-year growth rates, narrowed time period: 2001 Q2–2008 Q4

Dependent variable: Regressors:	**Exports**	**GDP**	Manufacturing exports	Manufacturing output
Constant	**7.261*** **	**0.549**	2.732	4.538*
	(2.132)	**(0.612)**	(2.620)	(2.313)
First lag of dependent variable	**0.454*** **	**1.112*** **	0.760***	0.428*
	(0.163)	**(0.243)**	(0.189)	(0.217)
Second lag of dependent variable		**-0.220**		
		(0.241)		
$REER_{t-1}$ (W_x and W)	**-0.675*** **	**-0.049**	-0.241	-0.335
	(0.186)	**(0.043)**	(0.292)	(0.244)
Adjusted R-squared	**0.500**	**0.765**	0.322	0.161
Number of observations	**31**	**31**	31	31
Durbin-Watson statistic	**1.709**	**1.786**	1.917	1.782

Notes: Standard errors are reported in parentheses. *, **, and *** denote significance at the 90, 95, and 99% levels. All variables are at constant prices, seasonally adjusted (except the REER) and stationary based on the ADF test. The real effective exchange rate is deflated by PPI and the weights are based on international trade volumes in SITC categories 5–8 in 2010 (source: Czech National Bank).

Table A7: Exchange Rate Elasticity Estimates on Macro Data (AR-X model)

All variables in year-on-year growth rates, full sample: 1997 Q2–2012 Q4

Dependent variable:	Exports	GDP	Manufacturing exports	Manufacturing output
Regressors:	1997 Q2–2012 Q3	1997 Q3–2012 Q3	2001 Q2–2012 Q4	2001 Q2–2012 Q4
Constant	2.620**	0.320**	1.170	0.966
	(0.862)	(0.229)	(1.190)	(0.969)
First lag of dependent variable	0.793***	1.548***	0.886***	0.815***
	(0.073)	(0.099)	(0.102)	(0.101)
Second lag of dependent variable		-0.648***		
		(0.099)		
$REER_{t-1}$ (W_x and W)	-0.443***	-0.028	-0.279	-0.208
	(0.111)	(0.023)	(0.223)	(0.182)
Adjusted R-squared	0.677	0.919	0.663	0.593
Number of observations	62	61	47	47
Durbin-Watson statistic	1.617	2.054	1.689	1.654

Notes: Standard errors are reported in parentheses. *, **, and *** denote significance at the 90, 95, and 99% levels. All variables are at constant prices, seasonally adjusted (except the REER) and stationary based on the ADF test. The real effective exchange rate is deflated by PPI and the weights are based on international trade volumes in SITC categories 5–8 in 2010 (source: Czech National Bank).

Table A8: Exchange Rate Elasticity Estimates on Macro Data (via equations 16 and 35)

All variables in year-on-year growth rates, various time periods

Dependent variable:	**GDP**	GDP	Manufacturing output	Manufacturing output
Regressors:	2001 Q2–2008 Q4	1997 Q3–2012 Q3	2001 Q2–2008 Q4	2001 Q2–2012 Q4
Constant (a_{00})	**0,397**	0,091	0,396	-1,537
	(0.438)	(0.174)	(2.341)	(1.007)
First lag of dependent variable (a_{01})	**0.637*****	1.287***	0,055	0.244*
	(0.131)	(0.121)	(0.245)	(0.134)
Second lag of dependent variable (a_{02})		-0.572***		
		(0.110)		
Exports to GDP ratio (a_1)	**0.148****	0.063*	0,113	0,296
	(0.060)	(0.034)	(0.375)	(0.270)
Imports to GDP ratio (a_2)	**-0,075**	-0,034	0,529	0.515*
	(0.061)	(0.041)	(0.360)	(0.282)
First lag of TFP (a_3)	**0.248***	0.225**	0,884	0.785**
	(0.139)	(0.098)	(0.630)	(0.362)
Implied ER elasticity of exports	**-0.752*****	-0.775***	-0,116	-0,215
($w_x = a_3 - 1$)	**(0.139)**	(0.098)	(0.630)	(0.362)
Implied ER elasticity of output	**0,110**	0.164*	0,777	0,529
($w = a_3 + \exp(-a_1) - 1$)	**(0.152)**	(0.098)	(0.807)	(0.415)
Adjusted R-squared	**0,838**	0,929	0,311	0,737
Number of observations	**31**	61	31	46
Durbin-Watson statistic	**1,968**	1,993	1,971	2,128

Notes: Standard errors are reported in parentheses. *, **, and *** denote significance at the 90, 95, and 99% levels. All variables are at constant prices, seasonally adjusted and stationary based on the ADF test. TFP is a European Commission estimate based on the production function approach (published annually; we use quarterly interpolation).

Table A9: Descriptive Statistics of the Macro-level Data

Year-on-year growth rates unless indicated otherwise

	Mean	Standard deviation	Min	Max	Number of observations	Observed period
*National Accounts:**						
GDP	2.5	3.0	-5.6	7.3	63	1997 Q1–2012 Q3
Exports	8.2	8.1	-20.5	22.4	63	1997 Q1–2012 Q3
Imports	6.8	7.6	-20.0	19.5	63	1997 Q1–2012 Q3
Exports to GDP (% y-o-y)	5.7	6.7	-16.7	23.4	63	1997 Q1–2012 Q3
Imports to GDP (% y-o-y)	4.3	6.1	-15.8	15.6	63	1997 Q1–2012 Q3
Exports to GDP (ratio)	0.60	0.15	0.34	0.88	67	1996 Q1–2012 Q3
Imports to GDP (ratio)	0.59	0.11	0.37	0.77	67	1996 Q1–2012 Q3
Exports to imports (ratio)	1.02	0.07	0.87	1.14	67	1996 Q1–2012 Q3
*Manufacturing:***						
Output	4.3	8.9	-23.9	15.9	48	2001 Q1–2012 Q4
Exports	8.4	10.9	-23.9	25.0	48	2001 Q1–2012 Q4
REER***	1.9	5.4	-8.0	13.5	64	1997 Q1–2012 Q4

Notes:

* Constant prices of 2005, seasonally adjusted. Data released in December 2012. Source: Czech Statistical Office.

** Index 2005=100. Output at constant prices. Export deflated by the export deflator from the National Accounts, seas. adj. by Tramo/Seats.

*** Index 2005=100. Deflated by PPI. Weights based on international trade volumes in SITC 5–8 in 2010. Source: Czech National Bank.