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Essays on Foreign Capital, External Finance and Trade of European Firms

Peter Tóth

Dissertation

Prague, October 2013

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Abstract

The dissertation is composed of three chapters studying empirical questions of corporate finance, European integration and international trade using firm-level data. The first chapter is focused at the macroeconomic and firm-specific determinants of attracting foreign investors to Czech firms before Czech EU membership. The second chapter deals with testing the effect of introducing the European single currency on the likelihood of equity and corporate debt issues by firms in the eurozone. The third chapter asks to what extent can cheaper imported inputs offset the loss in export sales due to domestic currency appreciation in Czech manufacturing firms around the EU accession period.

In the first chapter we study what factors attract foreign capital to Czech firms during the mid-transition, pre-EU-accession period 1997-2002. While the foreign owners' influence on firm performance has been widely studied, the same is not true for the opposite direction of causality. We consider macroeconomic, industry- and firm-level indicators of a firm's attractiveness to foreign capital. Using panel data techniques, we estimate linear models, limited dependent variable models and a hazard model with foreign ownership as the response variable. We find that foreign investors are likely to come from countries with higher corporate taxes and labor costs compared to the Czech Republic. As regards firm-level indicators, foreign investors favor larger firms, larger market shares and high concentration of ownership. Contrary to what we expected, indicators of the financial performance of a company did not seem to be significant attractors of foreign capital.

In the second chapter we test whether the introduction of the euro increased the likelihood of equity and corporate debt issues by firms in the eurozone. We hypothesize that the euro led to lower foreign exchange risks and transaction costs for international investors. This reduced the costs of issuing equity and corporate debt. Using a panel of about 6,000 Western European listed firms observed from 1995 to 2002, we estimate the likelihood of issuing equity or debt before and after eurozone entry and compared to firms with no "eurozone experience". At the same time, individual leverage targets, firm size and profitability are taken into account. We find a positive euro effect for issuing debt and external finance in general. A positive effect for issuing equity holds only in industries with heavy external finance dependence. Furthermore, the findings are consistent with other studies and capital structure theories suggesting that firms tend to revert to their leverage targets.

In the third chapter we ask to what extent can Czech exporters cushion the impact of currency appreciation shocks by using imported intermediates. A partial equilibrium

model with heterogeneous firms is applied. Producers can serve the domestic market, export final goods, or import inputs. In the model, an exogenous exchange rate shock simultaneously affects the variable costs and revenues associated with exports and imports. The impact of a hypothetical 1% appreciation of the domestic currency on sales is estimated using a panel of 7,356 Czech manufacturing firms observed from 2003 to 2006. The estimates are identified from within-firm variation in trade strategies, which is probably associated with the lifting of trade barriers due to Czech EU membership since 2004. For firms that both export and import, a drop in total sales of 0.2%, a drop in export sales of 0.8%, and a rise in domestic sales of 0.2% are predicted.

Abstrakt

Disertace se skládá ze tří kapitol zkoumajících empirické otázky firemních financí, evropské integrace a mezinárodního obchodu použitím firemních dat. V první kapitole se díváme na makroekonomické a firemně specifické determinanty přilákání zahraničních investorů do českých firem před českým členstvím v EU. V druhé kapitole testujeme efekt zavedení evropské společné měny na pravděpodobnost, že firmy z eurozóny emitují vlastní jmění a korporátní dluh. V třetí kapitole odhadujeme, do jaké míry můžou levnější dovezené vstupy kompenzovat ztrátu exportních tržeb v důsledku apreciace domácí měny z pohledu českých firem v zpracovatelském průmyslu v periodě kolem vstupu do EU.

V první kapitole zkoumáme, že které faktory přitahují zahraniční kapitál do českých firem v průběhu střední doby transformace, v periodě před vstupem do EU od 1997 do 2002. Zatímco vliv zahraničních vlastníků na výkon firem byl v literatuře často bádán, totéž neplatí pro opačný směr kauzality. Bereme v úvahu makroekonomické, odvětvově- a firemně specifické indikátory atraktivnosti firem pro zahraniční kapitál. Použitím technik panelových dat, odhadujeme lineární modely, modely s omezenými závislými proměnnými a hazardní model se zahraničním vlastnictvím jako odezvovou proměnnou. Nalézáme, že zahraniční investoři pocházejí s vyšší pravděpodobností ze zemí, kde jsou vyšší daně z příjmů firem a náklady práce v porovnaní s Českou republikou. Investoři také upřednostňují větší firmy, větší trhové podíly a vysokou koncentraci vlastnictví. Oproti očekáváním však indikátory finančního výkonu firem se nezdají být významnými atraktory zahraničního kapitálu.

V druhé kapitole testujeme, jestli zavedení eura zvýšilo pravděpodobnost, že firmy z eurozóny emitují vlastní jmění nebo korporátní dluh. Podle naší hypotézy, euro mohlo vést ke snížení kurzových rizik a transakčních nákladů pro mezinárodní investory. To snižovalo náklady pro emitování vlastního jmění a firemního dluhu. Použitím panelu přibližně 6,000 západoevropských firem obchodovaných na akciových trzích a pozorovaných mezi roky 1995 a 2002, odhadujeme pravděpodobnost emise vlastního jmění anebo dluhu před i po vstupu do eurozóny v porovnání s firmami bez „zkušenosti s eurozónou". Zároveň bereme v úvahu individuální cílové hodnoty finanční páky, velikost a ziskovost firem. Nalézáme kladný efekt eura na pravděpodobnost emitování dluhu a externích financí všeobecně. Kladný efekt v případě emise vlastního jmění se potvrzuje pouze v odvětvích s vyšší závislostí na externích financích. Dále, naše výsledky jsou konzistentní s jinými studiemi a teoriemi struktury kapitálu, které navrhují, že firmy mají tendenci se vracet

ke svým cílovým hodnotám finanční páky.

V třetí kapitole se ptáme, že do jaké míry můžou čeští exportéři tlumit dopad kurzových apreciačních šoků použitím dovezených meziproduktů. Aplikujeme model parciální rovnováhy s heterogenními firmami. Výrobci můžou obsluhovat domácí trh, exportovat hotové výrobky, anebo dovážet vstupy. Exogenní kurzové šoky v modelu simultánně ovlivňují variabilní náklady a tržby spojené s exporty a importy. Dopad hypotetické 1% apreciace domácí měny na tržby je odhadován pomocí panelu 7,356 českých firem ze zpracovatelského průmyslu pozorovaných od roku 2003 do 2006. Odhady jsou identifikovány díky změnám strategií exportování a importování v rámci firem, které jsou pravděpodobně spojeny s odbouráním bariér mezinárodního obchodu v důsledku českého členství v EU od roku 2004. Pro firmy, které exportují i importují, predikujeme pokles celkových tržeb o 0.2%, pokles exportních tržeb o 0.8% a růst domácích tržeb o 0.2%.

Introduction

The dissertation is composed of three chapters studying empirical questions of corporate finances, European integration and international trade using firm-level data. The first chapter tests the factors attracting foreign capital in Czech firms during the mid-transition, pre-EU-accession period from 1997 to 2002. The second chapter studies the effect of introducing the European single currency on the external financing of Western European firms before and after 1999. The third chapter deals with currency shocks to export sales of importers in the Czech manufactuding sectors around the Czech Republic's accession to the EU.

In the first chapter, co-authored with Petr Zemčík, we study the determinants of attracting foreign capital in Czech firms during the period 1997-2002. The effect of foreign ownership on firm performance in transition countries has been widely studied. The same is, however, not true for the opposite direction of causality. Which is, whether foreign capital targets the best performing firms in post-communist economies. Using a panel dataset of Czech firms with detailed information on the ownership structure and the nationalities of the owners, we relate various macroeconomic and firm-level indicators of a firm's attractiveness to the percentage stake of foreign owners. Our study extends the literature, firstly, thanks to the data on the nationality of the foreign owners. We use this information to construct macro-level indexes in the form of foreign stakes-weighted cross-country differentials, which vary firm-by-firm. Secondly, the panel dimension of the data allows us to control for time-invariant unobserved characteristics or to estimate a hazard model of acquiring a foreign investor. Thirdly, we apply a panel Granger causality

test to uncover potentially endogenous regressors and to treat them subsequently by some of the estimators.

Our empirical models cover three classes of estimators. First, a linear regression model is assumed and panel estimators are applied. The advantage of the first approach is that we can control for the time-invariant unobserved characteristics of the firms, treat the potential endogeneity of some of the indicators and use the percentage of foreign ownership as a continuous variable. Second, we estimate limited dependent variable models with fixed or random effects. Their advantage lies in their ability to model the switch between zero and non-zero foreign ownership stakes explicitly. However, the endogeneous regressors issue cannot be treated in this setup. The third approach applies Cox's proportional hazard model to estimate the probability of acquiring some foreign capital, provided the firm is purely domestically owned. In the hazard model, however, we cannot control for endogenous regressors and macro-level indicators.

We find that macroeconomic factors, such as the regional investor sentiment and cross-country differentials in labor costs and corporate taxes, have a significant impact on attracting foreign owners. Among the firm-specific characteristics, high ownership concentration is a significant attractor of foreign capital according to all models considered. Other firm- and industry-level factors with a positive impact include firm size, market share and industry dispersion of returns (risk). Contrary to our expectations, various financial ratios of a company were not significant. The results imply that foreign investors look for favorable macroeconomic conditions, larger firms, and aim to control a firm or an industry. However, prior financial performance is not of key importance for investors from abroad.

In the second chapter, co-authored with Karin Jõeveer, we test the effect of the euro introduction on the external financing of Western European firms. According to our hypothesis, the European single currency reduced the costs of external finance of firms in the eurozone. This is attributed to lower risks and transaction costs related to currency conversion from the point of view of international investors. We ask whether firms in the eurozone were issuing more debt and equity after 1999 compared to other Western European firms. Other studies have documented the euro effect in terms of reduced costs of equity finance, a vanishing home country bias in financial investments or increased corporate debt financing in the eurozone. But, to our knowledge, this question has not been studied on firm-level data.

To test the effect of the euro on debt and equity issues we follow Hovakimian, Opler,

and Titman (2001). The authors suggest that leverage is mean-reverting. In other words, equity and debt issues should respond to deviations from target leverage. In addition, other controls, such as profitability and firm size, are considered. We extend the mentioned setup by adding a dummy variable indicating eurozone membership, a proxy for external finance dependence and controlling for firm-specific fixed effects. Given firms' ambitions to meet leverage targets, we expect equity and debt issues to be more likely during eurozone membership. Firm-level data are obtained from the Amadeus database of Bureau van Dijk. The unbalanced panel of about 6,000 listed firms observed from 1995 to 2002 covers most of the eurozone, other EU countries, as well as Western European countries outside the EU.

The results are in line with our hypothesis on the euro effect. Namely, after eurozone entry, firms were more likely to issue debt and external finance in general compared to firms outside the currency area. The likelihood of issuing equity in the eurozone increased only if the firm belongs to industries with heavy external finance dependence. Our additional findings are consistent with other studies (Hovakimian, Opler, and Titman (2001)) and capital structure theories suggesting that firms tend to revert to their leverage targets. Furthermore, more profitable firms were more likely to issue debt and less likely to issue equity, holding their growth opportunities equal.

In the third chapter we ask to what extent can exporters cushion the impact of currency appreciation shocks by using imported intermediates. Over recent years, Czech manufacturing exporters have repeatedly caught the attention of the media during episodes of abrupt appreciation of the domestic currency, which, it is claimed, wipe out their profit margins. 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. To answer the above question, we apply a heterogeneous firms model, where firms can export and import. This allows us to estimate the impact of a hypothetical 1% appreciation of the domestic currency on sales for different exporting and importing strategies. The above has not been done in the context of heterogeneous firms models. The estimates are identified through exogenous variation in firms' trade strategies, using a panel of 7,356 manufacturing firms observed around the period of EU accession.

In our partial equilibrium model, firms are heterogeneous in their productivities and can serve the domestic market, export final goods, or import inputs. We introduce an exogenous exchange rate shock, which simultaneously affects the variable costs and revenues associated with exports and imports. The exchange rate elasticity estimates for

different trade strategies follow from the equilibrium sales equation implied by the model. Before the sales equation was estimated, we needed to fit total factor productivity from a production function with imported intermediates as productive inputs. In addition, to correct for the non-random selection of firms into exporting or importing, a multinomial probit model was estimated year-by-year. The fitted probabilities were then used to instrument the binary indicators of export and import. Finally, the parameter estimates of the sales equation allowed us to identify some of the structural parameters of the model as well as the exchange rate elasticities of sales.

The findings suggest that cheaper imported intermediates can partially offset the adverse effect of an appreciation shock to export sales. If a hypothetical appreciation of the domestic currency causes a drop in export sales of a non-importer by 1%, the same impact is only 0.8% for an importer of foreign inputs. The above is roughly in line with our benchmark macro-level estimates. Further results related to the estimation of firm-level production functions concur with other studies. First, we confirm the importance of measurement error correction in capital. Second, we find that imported intermediates have a positive effect on total factor productivity.

Chapter 1

What Makes Firms in Emerging Markets Attractive to Foreign Investors? Micro-Evidence from the Czech Republic[1]

Abstract

We use a panel of Czech firms to enhance existing literature where the dependent variable is foreign ownership. In our estimation, we control for endogeneity and unobserved effects using standard methods complemented by tests for heterogenous Granger-causality. We also model foreign ownership as a response variable in a hazard model and consider sorting by foreign owners rather then by domestic firms. We find that foreigners target firms with a greater ownership concentration in industries's with higher level of risk, in countries with lower labor costs and corporate income taxes.

JEL classification: G3, F21,C23-C25
Keywords: foreign ownership, endogeneity, causality, fixed effects, hazard model, truncated sample

[1]This paper has been published in Tóth, P. and P. Zemčík 2006. "What Makes Firms in Emerging Markets Attractive to Foreign Investors? Micro-Evidence from the Czech Republic." CERGE-EI Working Paper 294.

1.1 Introduction

The implications of the ownership structure in general have been well studied in emerging markets, and especially in the Czech environment, which provides an ideal natural experiment of a transition from 100% state ownership into a wide spectrum of ownership types. The effects of ownership concentration on performance controlling for various types of owners for (mostly newly privatized) firms have been investigated in Djankov (1999), Kočenda (2002), Lízal and Švejnar (2002), Lízal (2002), and Cull, Matesova, and Shirley (2002). The impact of the ownership structure on stock prices (i.e. the firm's value) has been considered for instance by Makhija and Spiro (2000) and Pajuste (2002).

The research studies including foreign ownership as one of the explanatory factors indicate that the presence of foreigners among owners improves performance of the firms. Lízal and Švejnar (2002) find that, based on a number of performance indicators, presence of foreign owners has a positive effect on the long-term performance of former state enterprises. Lízal (2002) demonstrates that foreign presence reduces the probability of bankruptcy and notes that employment in the foreign controlled enterprises has been growing in the late 1990's, and financial indicators have also been superior compared to those of firms with no foreign involvement. Hanousek, Kočenda, and Švejnar (2007) illustrate that the presence of majority foreign ownership has a positive effect on private Czech firms. Sabirianova-Peter, Švejnar, and Terrell (2012) find that foreign-owned firms occupy the top deciles of overall efficiency not only in the Czech Republic but also in Russia. As for stock prices, Makhija and Spiro (2000) conclude that ownership stakes of foreigners are positively related to share values.

Since foreign ownership clearly improves overall performance of Czech (and other) firms, it is of interest to investigate the motives of foreigners behind their investment in the Czech Republic and hence reverse causality of the studies cited above. The voucher privatization provided the first opportunity to investigate what causes a foreign investor to invest in a Czech company, and once she invests, what is the typical magnitude of ownership. Anderson, Jandík, and Makhija (2001) find that foreigners preferred safe and profitable firms, which did not have shares reserved for Czech insiders. Furthermore, the structure and size of the foreign investor's equity stake decreased with profitability and a high degree of tangible assets and increased with the level of indebtedness and the variability of industry profitability. By definition, the sample of data used in Anderson, Jandík, and Makhija (2001) consists of firms participating in the voucher privatization

process. The choice is given by the fact that some basic accounting data were made available to investors by the Center for Voucher Privatization of the Czech Ministry of Finance. The nature of the data provides basis for a simple cross-sectional analysis with only 41 foreign-owned firms. This dataset can be significantly enhanced due to the fact that high quality accounting data have been collected for a large number of firms, and the sample is not limited to the firms included in the first wave of privatization. The data have been recorded at annual frequency and we can make use of panel data techniques.

We use a panel data of Czech firms, most of which are not listed on the Prague Stock of Exchange. We relate foreign ownership in a firm to various macroeconomic and firm specific variables. The macroeconomic variables include investor sentiment,[2] corporate income tax, labor costs, and labor intensity differentials between the Czech Republic and the home country of foreign stakeholders (see Kulawczuk, Bąk, and Szcześniak (2005)). The firm-level characteristics contain various measures of profitability, risk, size, ownership concentration, market share, etc. The variable of interest is the foreign stake in a given firm. While previous studies (e.g. Sabirianova-Peter, Švejnar, and Terrell (2012)) only used a dummy indicating foreign ownership greater than 50%, we are able to extract more detailed information regarding foreign stakeholders, namely their percentage stakes in the company and countries of origin.

This additional information together with panel structure of our dataset helps us to extend research where the dependent variable is foreign ownership in a number of ways. We test for Granger causality of the dependent variable and potentially endogenous explanatory variables. Our candidates for endogenous variables are ownership concentration and the industry's volatility of the accounting rate of return (our measure of systemic risk). We use new panel data tests for heterogenous Granger-causality proposed in Dumitrescu and Hurlin (2012) which are suitable for data with a short time-series dimension. We find that foreign ownership Granger causes both variables and vice versa which is consistent with our prior of their endogeneity. In our subsequent estimation, we control for this endogeneity and account for unobserved effects, neither of which was a part of the analysis in Anderson, Jandík, and Makhija (2001).

For the sake of robustness, we employ several econometric techniques in our estimation of the impact of macro and firm-level variables. Namely, we estimate a linear model by OLS as a benchmark, use fixed effects and first differences transformations, and then apply

[2]Sokalska (2001) defines the international investor sentiment as an unobservable variable approximating joint movements of equity markets in Poland, Hungary, and the Czech Republic.

2SLS. In addition, we consider panel data versions of limited dependent variable probit and logit models and a corner solution panel data Tobit. Based on some specification tests and taking into account limitations of our data, we center our attention on results from fixed-effects-2SLS and first-differences linear models. However, our estimates are mostly robust to the use of a particular estimation strategy.

Our results indicate that macro variables such as international investor sentiment and employment compensation, labor costs and corporate tax differentials all have a positive impact on stakes of foreigners. The strong effect of the differentials echoes the notion that profitability of marginal investment by foreign firms increases by investing abroad (see Kulawczuk, Bąk, and Szcześniak (2005)). From firm specific variables, increases in the standard deviation of a firm's industry's rate of return, ownership concentration, size and industry share imply an increase in foreign ownership.

We further examine our data using different prospectives. We start with an attempt to answer a question of what makes domestic firms to be a more likely target of foreigners. We use a proportional hazard model to estimate probability of a switch from a zero foreign ownership to a positive one. Foreign ownership is hence viewed in the terms of duration, similarly to literature on unemployment. Due to data limitations, we assume single-spell data and time-invariant covariates. Conditional on a foreign investor purchasing a share in a Czech firm, foreign investors still prefer large firms with high ownership concentration, in an industry with a greater variability of profitability.

Finally, we consider a foreign owner (rather than a foreign owned firm) to be our sorting criterion for a dependent variable. We thus form a truncated sample of stakes of foreign investors with corresponding explanatory variables. The sample is truncated because we do not possess any information regarding foreign firms which have decided not to purchase a company share of a firm in the Czech Republic. The pooled OLS again show significance of coefficients on the industry systemic risk and ownership concentration. Interestingly, the impact of macro variables is limited in this case.

The remainder of the paper is organized as follows. Section 1.2 motivates used explanatory variables, Section 1.3 describes our data and their source, Section 1.4 characterizes the employed econometric techniques and analyzes the results, and Section 1.5 concludes.

1.2 Determinants of Foreign Ownership

In this section, we characterize general features of a firm, which might make it more attractive to foreign investors. We divide these characteristics into two groups, firm-level variables and macro-level variables, respectively. We formulate our predictions based on a combination of theory, common sense, and previous empirical observations. We adopt a perspective of a foreign firm, which is trying to maximize Net Present Value of its investments, some of which may take place abroad. Many of the variables are inspired by the literature on portfolio investment. As many of the firms in our data are not listed on any stock exchange, we employ various analogies of the standard stock market related variables. Our predictions will be tested on Czech data using cross-sectional and panel data estimation methods in Section 1.4.

1.2.1 Firm-level Determinants

Profitability

Profitable firms tend to be more attractive and hence preferred by foreign investors. There are several potential measures of profitability. We use the accounting rate of return (ARR) defined as the annual average of net income to book value of shareholder's equity. An alternative for firms listed on the stock exchange is the B/M ratio, defined as the book value of equity divided by the market value of equity at the end of the fiscal year. Typically, growth firms are associated with a low B/M ratio while the value firms have a high B/M ratio. Using the results in Fama and French (1995), B/M can be substituted by ARR for firms not listed on any stock exchange.

Risk

Risk considerations are somewhat more complex. To compensate for various costs related to cross-border investment, a foreign investor might be willing to accept a higher level of risk provided it is priced properly, i.e. higher risk is associated with the higher rate of return. Therefore, foreign investors prefer firms with lower idiosyncratic and higher systemic risk. The standard stock market measures of systemic and idiosyncratic risks are based on the estimation of the market model, in which a firm's daily excess return is regressed on the daily excess return on the market portfolio. The estimated coefficient is the measure of systemic risk beta and the residual variance from the regression is the measure of idiosyncratic risk. However, these measures cannot be used in our case

as only a very small number of firms is listed on the Prague Stock Exchange. One possible proxy for systemic risk of non-listed firms is the intra-industry variance of ARR. The idiosyncratic risk for individual firms is well characterized by the data on capital structure, which describe the financial health of the company. Long term financial distress is captured for example by the ratio of total liabilities to total assets (the leverage ratio). The short term financial distress is reflected in the current ratio i.e. the ratio of current assets to current liabilities at the end of a fiscal year.A company's solvency is characterized by the ratio of the cash-flow to the market value of equity (or the book value).

Firm Size

Contrary to predictions of standard finance models such as the capital asset pricing model, size explains a large portion of both cross-sectional and time-series variation in returns (see Fama and French (1992, 1993) for the summary of existing literature and extensive empirical analysis). Moreover, greater size may be cause firms to be well-known abroad, which in turn stimulates foreign investment - see Kang and Stultz (1997), Dahlquist and Robertsson (2001), and Lin and Shiu (2003). The size is significant in all these studies, even after accounting for other firm's characteristics. Also, large firms are more likely to be successful in utilizing the expertise and capital of foreign investors. Hence, we are likely to observe that foreign investors target larger firms. We measure size as the total book value of firms' assets.

Ownership Concentration

In an environment, where minority owners are not well protected and where acquiring a portion of a domestic company requires a lot of investment in terms of opportunity costs, foreigners are likely to target a controlling share of the company i.e. greater than 50%.[3]

Market Share

It is also likely that foreigners tend to purchase domestic firms to get access to a local market. In this case, they tend to prefer local firms with higher market shares.

Other Factors

Here we include variables, which are difficult to fit into other categories, such as staff costs per sales, value added per sales, etc. Staff costs per sales are likely to play a role for firms from developed countries investing in emerging economies. We will discuss this issue

[3]In the Czech Republic, this relationship is likely to be even stronger in the near future. In May 2005, §813 of the Czech Business Code was amended to allow forced buy-outs of minority shareholders by shareholders owning 95% and more of a given company.

below in some detail when we comment on the role of average employee compensation. Value added is typically going to be high in industries with a qualified labor force. For example, a German firm may buy a controlling share in a Czech firm in the same industry when it can expect lower labor costs for relatively highly qualified labor.

While there are other candidates for the firm-level factors (e.g. high export ratios), we do not mention them here due to restricted data availability.

1.2.2 Macro-level Determinants

International Investor Sentiment

This variable is inspired by Sokalska (2001) who defines the international investor sentiment as an estimated state variable from a Markov chain model. The Markov chain is estimated using data on equity returns fron the Czech Republic, Hungary, and Poland. The selection of the three countries is given partly by availability of the data and partly by the fact that they are considered leaders in the region. The stock market indices from the stock exchanges in Prague, Budapest, and Warsaw move closely together. The impact of the local macroeconomic fundamentals is rather small, perhaps due to small capitalization, volume of trading, and liquidity. Sokalska (2001) documents that the estimated process is correlated with global macroeconomic fundamentals, from the US term structure to emerging markets indices.

While one can only speculate on the relative importance of this latent variable as compared to the firm level determinants of foreign ownership, the following hypothesis can be formulated: The more favorable the global investment climate, the higher foreign ownership. To avoid issues connected with selection of a correct specification for the Markov chain process, we construct a simple variable based on monthly movements of the three stock markets. First, we define a dummy for each market, with the dummy being equal to one if the stock market in question is above its long-term average on a given month. The average of monthly sums in a given year gives us a desired proxy for the international investor sentiment.

Corporate Income Tax and Labor Costs

Foreign investors may be interested in buying a share in domestic enterprises if they can save either on taxes or on labor costs as compared to their home countries. Kulawczuk, Bąk, and Szcześniak (2005) formalize this idea and construct an index, which characterizes

additional profitability of a firm investing abroad. The index is based on the assumption of the foreign investor being able to achieve the same level of productivity in a targeted company as in her base country and considers differences in both corporate income taxes and labor costs between two countries, the home country of the foreign investor and the country where she invests.

We split the two effects and analyze them separately. Our dataset of Czech firms allows us to identify a country of origin for foreign owners. Using regression analysis, we estimate the effects of corporate income tax and labor cost differentials. We use statutory corporate income taxes and various measures of labor costs and also include labor intensity of the base country.

1.3 Data

We use an unbalanced panel of Czech firms from all economic sectors. The data was collected by Aspekt s.r.o. starting from 1993. The largest part of our sample consists of firms in manufacturing and trade. The sample period starts in 1997 and ends in 2002, with only a few observations in 2002. The maximum number of annual observations is 1979 firms in year 2000. We restrict our sample to 1997 onwards due to the availability of some of the macro-level variables and because the number of firms is small before 1997.

The variables are defined in Appendix A. Descriptive statistics for the firm-level variables are respectively given in Table 1.1 for the whole panel and in Tables 1.2, 1.3, and 1.4 for cross sectional data by years. Several patterns emerge. From 1997 until 2002, the size of firms in terms of total assets increased with annual growth rates exceeding inflation. To eliminate the effect of inflation, we deflate tangible assets using CPI from the Czech Statistical Office (base year 2000). Profitability measured by the accounting rate of return decreased over the observed time period, from 24.29% to 8.13%. Figure 1.1 shows the histogram in percentages. Lower profitability may be a sign of a maturing market in the Czech Republic and the tendency of the Czech firms to invest in projects with a lower level of risk (approximated by the standard deviation of the industries' accounting rates of return, *indsarr*). Indebtness (see variable *lever*) decreased as well, suggesting that firms relied more on internal resources for growth. Current ratio and cash-flows per sales vary from firm to firm, and over time, the solvency rate decreased over the considered time-period. Staff costs per sales showed a steady upward trend, as opposed to th drop

in value added per sales towards the end of the observed period. The Herfindahl index of ownership concentration was increasing steadily during most of the sample and stabilized above 6000 in the last two years, 2001 and 2002.

The variable of interest is the share of registered capital owned by foreigners. Studies on Czech firms using foreign ownership as a dependent variable have only been able to distinguish between foreign ownership greater or smaller than 50% (e.g. Sabirianova-Peter, Švejnar, and Terrell (2012)). The distinction is made at the level of the Czech Statistical Office. However, we are able to enhance this information in several ways. Using the Aspekt database, we determine the exact percentage of each firm owned by foreigners and their countries of origin. We then track changes in the percentage through time. We are not aware of any study, which uses a finer description of foreign ownership in emerging markets. The descriptive statistics for the whole sample and by years are reported in Tables 1.1, 1.2, 1.3, and 1.4, respectively. The level of foreign ownership had been increasing steadily since 1997 (8.53%) until 2002 (12.42%). Figure 1.2 shows the histogram of foreign ownership for the whole sample. We can see that most of the data points are concentrated at 0 and then at 100%. The distribution of only positive values is displayed in Figure 1.3. The more detailed view indicates higher accumulation of foreign ownership for values greater than 50% i.e. stakes allowing to control a given company.

The group of macro-level explanatory variables consists of the investor sentiment, labor costs per value added in manufacturing, the employee compensation rate, and corporate income taxes. The investor sentiment (Table 1.6) greater than 1.5 indicates that regional stock markets have been mostly increasing during a given year (see Table 1.5). Labor costs per value added in manufacturing are typically lower in new member countries of the European Union (Table 1.7). The employee overall compensation rate is higher in most countries investing in the Czech Republic (exceptions are either new members of the EU or less developed countries). Finally, corporate income tax rates in Table 1.8 are on average lower in new member countries, and there is an overall tendency to decrease the rates.

1.4 Estimation

In this section, we investigate the determinants of foreign ownership from several perspectives. First, to examine interaction between the dependent variable and (potentially)

endogenous explanatory variables, we test for causality using a technique suitable for panel data with a short time span. Second, we estimate a linear model (LM) with the stakes of foreigners as a dependent variable. We account for unobserved effects and endogeneity by first differences (FD) and fixed effects (FE) models in combination with 2SLS. Then we employ limited dependent variable probit and logit panel data models complemented by a corner response Tobit panel data model. Third, we use a hazard model to find what makes domestic firms likely to be targeted by foreigners. Finally, we estimate a truncated regression model where the focus is on foreign owners rather then foreign owned firms.

Causality

First, we would like to investigate the direction of causality between the dependent and two candidates for endogenous explanatory variables, respectively the ownership concentration and the standard deviation of the accounting rate of return in an industry. The high ownership concentration makes a domestic firm an easier target since a foreign investor needs to negotiate with a smaller number of owners. The variability of rate of return is almost surely affected by the entrance of foreigners, and it is possible that high systemic risk industries are targeted more often since the variability of the foreign acquirerer's profit is likely to be lower. While causality and endogeneity are obviously not identical concepts, evidence of causality is symptomatic for the presence of endogeneity between stakes of foreigners and the two above-mentioned variables. A suitable test for Granger causality in panel data with a short time-series dimension is proposed in Dumitrescu and Hurlin (2012)[4].

Consider two stationary variables, y and x_j, and the following linear model:

$$y_{it} = a_i + \sum_{l=1}^{L} \gamma_i^{(l)} y_{i,t-l} + \sum_{l=1}^{L} \delta_i^{(l)} x_{i,j,t-l} + \epsilon_{it}, \tag{1.1}$$

where e_{it} are normally i.i.d. with zero mean and finite heterogeneous variances and $\epsilon_i = (\epsilon_{i1}, ..., \epsilon_{iT})'$ are independently distributed across groups. The null hypothesis of the Homogeneous Non Causality (HNC) is:

$$H_0 : \ \delta_i = 0, \ \ \forall i = 1, ..., N, \tag{1.2}$$

[4]For an earlier application of the mentioned panel Granger causality test see Hurlin and Venet (2004)

where $\delta_i = (\delta_i^{(1)}, ..., \delta_i^{(L)})'$. The alternative hypothesis allows for N_1 individual units with no causality and for heterogeneous causality among the rest of the units. It is defined as:

$$\begin{aligned} H_1: \quad & \delta_i = 0, \quad \forall i = 1, ..., N_1, \\ & \delta_i \neq 0, \quad \forall i = N_1 + 1, ..., N, \end{aligned} \tag{1.3}$$

where $N_1 \in [0, N)$ is not known. Let W_{it} denote the Wald statistic associated with the individual test of H_0 for each $i = 1, \ldots, N$, and let $W_{NT}^{HNC} = (1/N) \sum_{i=1}^{N} W_{iT}$. Hurlin (2004) proves that for a fixed $T > 5 + 2L$, the approximated standardized statistic

$$Z_{NT}^{HNC} = \sqrt{\frac{N}{2 \times L} \times \frac{(T - 2L - 5)}{(T - L - 3)}} \times \left[\frac{(T - 2L - 3)}{(T - 2L - 1)} W_{NT}^{HNC} - L \right] \tag{1.4}$$

converges in distribution to $N(0, 1)$ as $N \to \infty$.

In the context of our research problem, y is foreign ownership and x_j is an explanatory variable of interest. We first verify stationarity by a panel data unit root test suggested in Im, Pesaran, and Shin (2003) and then investigate causality in both directions using the statistic (1.4), which can be easily applied to non-balanced panels. We only consider the lag $L = 1$ to meet the requirement of at least 8 observations, which we achieve by considering a longer time span than 1997-2002. The number of firms is not an issue in this case as we need 8 subsequent observations for each firm at any starting in any year from 1993 to 1996.

Evidence from panel data unit root tests, according the Z_{t-bar} statistic from Im, Pesaran, and Shin (2003), is somewhat mixed and depends on the inclusion of a time trend. For example, the null hypothesis of a unit root is rejected for $fshare$ when trend is not included (see Table 1.9), and it is rejected otherwise. The results are reversed for $herfin$. Since both of these variables are restricted with respective intervals from 0 to 1, and from 0 to 10,000 we view the two variables as stationary. For $insdarr$, the null is strongly rejected with no trend included and almost rejected at 10% level of significance, and hence, we consider it stationary as well. The results of causality tests are reported in Table 1.10. Foreign ownership is Granger-caused by the Herfindahl index and vice versa. We get similar results for foreign ownership and the industries' standard deviation of ARR. Presence of Granger causality indicates endogeneity among the investigated variables.

Foreign Stakes: A Standard Approach

We model the foreign ownership as the dependent variable (y_{it}) and the firm and macro characteristics as the explanatory variables (x_{it}'s), respectively. The estimated regression model then is:

$$y_{it} = \beta_0 + \beta_1 t + \beta_2 x_{i1t} + \ldots + \beta_k x_{ikt} + v_{it}, \quad i = 1, \ldots, n, t = 1, \ldots, k, \tag{1.5}$$

where

$$v_{it} = a_i + u_{it}. \tag{1.6}$$

We take logs for explanatory variables in the Czech currency such as total assets and relate others to sales when they can be negative to increase the number of observations (e.g. cash flows). The estimation of coefficients in (1.5) involves two challenges - endogeneity of some variables and treatment of the unobserved effect a_i. To handle endogeneity, we use 2SLS. The Herfindahl index and the industries' standard deviation of ARR are our endogenous variables. As instruments we use exogenous variables and lagged endogenous variables. We consider several potential properties of the error term v_{it} and use corresponding estimation methods. If a_i is correlated with some of x's, one can employ fixed effects or first-difference estimators, and if a_i is uncorrelated with explanatory variables, we can use the random effects estimator. In effect, we are using the 2SLS generalizations of simple panel data methods in the case of exogenous variables.

The Hausman specification test (see Table 1.11) rejects the the null hypothesis of the random effects model, and hence, we only report the results of the estimation of fixed effects models (LM-FE and LM-FE-2SLS) with a pooled OLS estimation of the LM as a reference (LM-OLS) - see Table 1.12. The Hausman test also rejects LM-FE in favor of the alternative LM-FE-2SLS. Since both models suffer from some minor but significant residual autocorrelation, we also estimate LM in first differences with results in Table 1.13 (LM-FD and LM-FD-2SLS).The price for robustness with respect to autocorrelation is a smaller number of observations. Here the Hausman test cannot reject H_0 : LM-FD and hence, we view LM-FD as preferable to LM-FD-2SLS. To summarize, our discussion of results will rely on the estimates of LM-FE-2SLS and LM-FD though there seems to be only minor quantitative differences across the used methods.

The macro variables are all significant and positive as hypothesized with the exception of the corporate tax differential in LM-FD, which is significantly negative, contrary

to expectations. However, if we look at the cross-country comparison of tax rates in Table 1.8, there is little year-to-year variation in the corporate tax differentials. Hence, the first difference estimator might be too harsh on the tax rate data compared to fixed effects.

Among the firm level variables, coefficients on the volatility of the accounting rate of return and the Herfindahl index of concentration are positive and significant in both LM-FE-2SLS and LM-FD. The positive coefficient for the industry's volatility supports the view that foreigners tend to focus on the industries with a high level of risk. This view is also consistent with our previous tests since high $insdarr$ Granger causes $fshare$.[5] The estimate of the coefficient for the index of concentration suggests that foreigners target firms with higher ownership concentration perhaps because they are easier to acquire.[6] Our estimation of hazard models below should shed more light on the mutual relationship between foreign ownership on the one hand and volatility of profits and ownership concentration on the other. In the LM-FE-2SLS, estimates of coefficients for solvency, real assets, and industry's share are significant. The estimate is negative for solvency and positive for the other two variables. While these estimates are not significant in LM-FD, they have the same sign as in LM-FE-2SLS. The estimate for the solvency coefficient is too small to be economically meaningful, but the other two estimates provide some evidence of foreigners purchasing bigger firms with high industry shares.

While the majority of values predicted by the linear model for foreign share $\hat{y}_{it}$ lie within the range of 0 and 1 with a higher frequency at 0, there are values lower than 0 and greater than 1, leading us to consider other ways of modelling the limited dependent variable as a robustness check of our results from the LM estimation. Note that in the LM setup we faced a censored dependent variable issue and left it untreated for simplicity. Therefore we expect the LM estimates in Tables 1.12 and 1.13 to be potentially biased towards zero[7].

Based on histogram 1.2, we can model the dependent variable as discrete i.e. $y_{it} = 0$ for a firm i without a foreign owner at time t and $y_{it} = 1$ otherwise. $y_{it} = 0$ if $y^*_{it} > 0$ and

[5]Obviously, the $fshare$ also has a consequent impact on $insdarr$.

[6]As was the case with the industry's standard deviation of ARR, the concentration changes after the foreigner purchases a stake in the company.

[7]Given that we are more interested to know the signs of the estimates and whether they are significantly different from zero, the results above can be viewed as conservative.

$y_{it} = 1$ if $y_{it}^* \leq 0$ where

$$y_{it}^* = \beta_0 + \beta_1 t + \beta_2 x_{i1t} + \ldots + \beta_k x_{ikt} + v_{it} = x_{it}'\beta + v_{it}, = \quad i = 1, \ldots, N, t = 1, \ldots, T, \quad (1.7)$$

with v_{it} defined in (1.6). For the fixed effect model,

$$\Pr[y_{it} = 1] = \Pr[y_{it}^* > 0] = \Pr[u_{it} > -x_{it}'\beta - a_i] = F(x_{it}\beta + a_i). \quad (1.8)$$

The density function is logistic in the case of the logit model and normal in the case of the probit model. The presence of a_i complicates the estimation; the computational burden can be reduced by obtaining the conditional likelihood for the logit model. A similar reduction is not possible for the normal density, and hence, it is standard to estimate only the random effects probit model (see Baltagi (2005)). The conditional fixed effects logit (Lo-FE) and random effects probit (Pro-RE) estimates are reported in Table 1.14. Assuming again that our dependent variable is continuous, we can use the Tobit model, which is well-suited for corner-solution responses (foreign ownership=0). The fixed effect Tobit is not estimated due to problems similar to those in the probit estimation. Therefore, we report only random effects Tobit (To-RE) estimates, also in Table 1.14. There are some shortcomings to these models. None of them account for endogeneity, which would make our estimation very complex and sensitive to various assumptions. Also, Pro-RE and T-RE ignore potential correlation between a_i's and explanatory variables. However, looking at the results from the linear models and comparing the LM-FE and LM-FD estimates with their 2SLS versions (Tables 1.12 and 1.13), we expect that most of the estimates not corrected for endogeneity are potentially downward biased[8].

In spite of these issues, conclusions based on the estimation of probit, logit, and Tobit models are roughly consistent with our results based on the LM. The coeficients estimates for macro variables have the expected sign when they are significant. The index of concentration and industry's share have significantly positive coefficients in all the regression models. Solvency's coefficient is either 0 or significantly negative as before, and the coefficient of industry's level of risk has a plus sign in all models and is significant in two cases.

Firms: A Hazard Model

[8]Given that we are more interested to know the signs of the estimates and whether they are significantly different from zero, the results above can be viewed as conservative.

Here, we treat foreign ownership as a response variable in the form of duration. We follow a firm and record when there is a change of foreign ownership state from 0 to positive[9]. Then we estimate the probability of this change conditional on firm characteristics. Since the time dimension of our data is limited and time-varying models typically require strict exogeneity of variables, we focus on a model with a hazard function conditional on time-invariant covariates. We also assume that we have only single-spell data[10].

Let S be the period during which the change occurs. The conditional hazard function is

$$\lambda(t; \mathbf{x}) = \lim_{h \downarrow 0} \frac{(t \leq S < t + h | S \geq t, \mathbf{x})}{h} \tag{1.9}$$

where $\mathbf{x}$ is the vector of explanatory variables. We assume that

$$\lambda(t; \mathbf{x}) = \kappa(\mathbf{x}) \lambda_0(t), \tag{1.10}$$

where $\kappa(\mathbf{x}) > 0$ and $\lambda_0(t) > 0$ is the baseline hazard. This is a proportional hazard model. $\kappa(\mathbf{x}) = \exp(\mathbf{x}\beta)$. We estimate β using the Cox (1972) approach. Only firm-specific variables are considered and we calculate standard errors robust to clustering on foreign ownership. Note that the advantage of Cox's proportional hazard model lies in parsimony, which is leaving the functional form of the baseline hazard function unspecified. This implies, though, that the model does not directly predict the instant probability and the expected duration of the spell. However, testing the covariates driving the hazard event is still possible. Therefore Cox's approach is well aligned with the goals of the present paper[11].

Results are reported in Table 1.15. Conditional on a decision to acquire a firm in the Czech Republic, foreigners seem to target larger firms with a higher ownership concentration and size in industries with a higher volatility of profits. In other words, the probability of a firm becoming owned at least partially by foreigners increases with the significant variables. The results support the view that foreign owners target firms with a higher level of risk. While their entrance can also increase the volatility by entering the

[9] As a sensitivity check of the results, we excluded foreign stakes smaller than 90 percent from the hazard indicator. The resulting estimates were very similar to the baseline case. Moreover, two additional explanatory variables became statistically significant. In order to draw conservative conclusions, we decided to keep the hazard indicator of zero-to-positive foreign ownership as the baseline case.

[10] As multiple spells may be partially associated with measurement errors, we found it safer to exclude those observations.

[11] It still remains a question whether the data support the proportionality assumption of Cox's model. To check this assumption, we performed a common proportionality test based on Schoenfeld residuals, as available in Stata. We were not able to reject proportionality at common significance levels.

Czech market, we cannot evaluate this hypothesis in the used hazard model. Foreigners also prefer firms with a higher ownership concentration, and we again cannot comment on whether their entrance increases the concentration or not.

Foreign Stakeholders

Here we attempt to find out what makes investors invest abroad. Data units are now selected by foreign owners. The dependent variable is a stake of a foreigner. The explanatory variables are macro variables and characteristics of the firm in which she owns a share. Obviously, we only have a truncated sample of foreign owners since we do not have any information for foreigners who did not invest abroad. We only use pooled OLS due to a small number of observations and technical difficulties connected with robustness to endogeneity, and unobserved fixed effects in panel data truncated regression models. Therefore, our estimates in Table 1.16 have to be treated with some caution. The industry's variability of profit and ownership concentration are again significant.

1.5 Conclusion

We investigate the determinants of foreign ownership using panel data on firms in the Czech Republic. We first investigate the mutual relationship between foreign ownership and two potentially endogenous variables: the Herfindahl index of concentration and the variability of industry's profit. Foreign ownership helps to predict the two variables and vice versa. Granger-causality in both directions can be interpreted as an indication of endogeneity. We proceed to evaluate the impact of various macro- and firm-level variables on the foreign share in a firm using a variety of econometric methods. We account for both endogeneity and unobservable fixed effects and employ a number of limited dependent variable and corner solution models. Share of foreigners in Czech firms are affected both by macroeconomic and firm-level factors. The macroeconomic factors with a positive impact on foreign ownership include international investor sentiment, differentials in labor costs, employee compensation, and corporate taxes across countries. On the other hand, important firm characteristics are variability of profit in the firm's industry, size, ownership concentration, and industry share.

The next step in our empirical investigation is an adoption of a novel perspective on how to study the effects of various factors on shares of foreigners in domestic firms.

We follow each firm and record the moment when it becomes at least partially owned by foreigners. By re-organizing the data in this manner, we can use the Cox hazard model to estimate the probability of the firm being foreign owned conditional on the foreigner's decision to purchase a Czech firm. Attractive targets are large in size, with high ownership concentration and industry risk. Yet another novel way to use our data is to sort foreign ownership percentages by foreign owners and not by firms. The truncated sample of

foreigners then can be used to quantify the effects of our macro and micro factors. Again, the factors with strong impact are the Herfindahl index of concentration and the standard deviation of the industry's rate of return.

1.6 Appendix: Data Definitions, Tables and Figures

A.1. Firm-level Variables:

logrta - logarithm of real total assets

arrps- accounting rate of return (operating profits over registered capital) per sales

lever - leverage: liabilities over total assets

current - current ratio: current assets over current liabilities

cashflps - cash-flow (or annual change in cash stock if the first is not available) per sales

solvency - solvency rate: cash-flow to registered capital

scps - staff costs per sales

vadps - value added per sales

insdarr - industry standard deviation of ARR

indshar - firm's share in industry sales

fshare - foreigners' share on registered capital

herfind - Herfindahl index of ownership concentration

quoted - dummy equals 1 if the firm's shares are registered with the Czech Securities Exchange Commission (SEC)

A.2. Macro-level Variables:

sentim - regional investor sentiment defined as the previous year's average of the monthly sums of three dummy variables, each of the dummies equaling to 1 if the monthly return on the Czech, Polish or Hungarian stock index respectively goes above its long-term average monthly return.

sentimq - regional investor sentiment for firms with registered shares at the Czech SEC, i.e. sentim multiplied by the dummy quoted. The dummy is based on information from the Aspekt database.

linta - labor intensity in a stakeholder's home country, based on the median firm's labor costs per value added ratio in a country's NACE D sector (manufacturing). For each firm we take the stake-size-weighted average of the ratios by different stakeholders. Data come from Eurostat.

ecra - average employee compensation. Compensation of employees has two main components: (a) Wages and salaries payable in cash or in kind; (b) The value of the social contributions payable by employers: these may be actual social contributions payable by employers to Social Security schemes or to private funded social insurance schemes to secure social benefits for their employees; or imputed social contributions by employers

providing unfunded social benefits. in a stakeholder's home country per average employee compensation in the Czech Republic. For each firm we take the stake-size-weighted average of the ratios by different stakeholders. Data are taken from OECD and Eurostat databases. *dcita* - statutory corporate income tax (CIT) rate in the stakeholder's home country minus the Czech CIT rate. For each firm we take the stake-size-weighted average of the differences by different stakeholders. The data source is Eurostat.

Table 1.1: Descriptive Statistics of the Firm Level Variables (1997-2002)

Variable	Obs	Mean	Std. Dev.	Min	Max
logrta	12470	11.21	1.33	2.95	17.37
arrps	12460	0.00	0.03	-2.82	1.06
lever	12470	0.64	0.72	-0.28	59.35
current	12470	5.25	83.09	0.00	5998
cashfps	12460	-3.31	308.89	-33020	1496
solvency	12470	1.13	41.86	-1764	1248
scps	12460	0.34	6.88	0.00	481
vadps	12460	-1.31	117.97	-12402	1
insdarr	12470	39.94	19.09	0.02	124.75
indshar	12470	0.00	0.02	0.00	1.00
fshare %	12470	9.85	28.36	0	100
herfind	12470	5885	3572	1	10000

Table 1.2: Descriptive Statistics of the Firm Level Variables, Years 1997-1998

1997

Variable	Obs	Mean	Std. dev.	Min	Max
logrta	2204	11.05	1.15	7.16	16.02
arrps	2202	0.00	0.02	-0.25	1.06
lever	2204	0.72	0.35	0.00	5.08
current	2204	3.03	20.43	0.00	686.97
cashfps	2202	0.86	31.88	-69.43	1361
solvency	2204	2.54	43.12	-313.32	1024
scps	2202	0.30	6.11	0.00	286.14
vadps	2202	0.02	4.95	-167.93	1.00
insdarr	2204	36.25	8.87	0.27	45.34
indshar	2204	0.00	0.00	0.00	0.11
fshare %	2204	8.53	26.53	0	100
herfind	2204	5392	3587	1	10000

1998

Variable	Obs	Mean	Std. dev.	Min	Max
logrta	2338	10.97	1.24	2.95	15.39
arrps	2336	0.00	0.06	-2.82	0.06
lever	2338	0.70	0.65	-0.28	18.83
current	2338	6.03	68.88	0.00	1970
cashfps	2336	-3.60	185.97	-8966	618.18
solvency	2338	0.99	34.04	-601.75	311.25
scps	2336	0.18	0.66	0.00	24.67
vadps	2336	0.05	4.37	-192.08	0.99
insdarr	2338	34.20	8.02	1.96	45.79
indshar	2338	0.00	0.00	0.00	0.11
fshare %	2338	9.22	27.61	0	100
herfind	2338	5796	3613	1	10000

Table 1.3: Descriptive Statistics of the Firm Level Variables, Years 1999-2000

1999

Variable	Obs	Mean	Std. dev.	Min	Max
logrta	2757	11.16	1.35	5.17	17.07
arrps	2756	0.00	0.01	-0.47	0.07
lever	2757	0.66	1.21	-0.01	59.35
current	2757	8.55	160.63	0.01	5998
cashfps	2756	-12.07	629.16	-33020	460.41
solvency	2757	1.57	51.01	-1422	1248
scps	2756	0.26	3.48	0.00	182.00
vadps	2756	-1.50	83.75	-4385	1.00
insdarr	2757	62.40	22.82	0.02	107.37
indshar	2757	0.00	0.02	0.00	0.69
fshare %	2757	9.37	27.74	0	100
herfind	2757	5900	3580	1	10000

2000

Variable	Obs	Mean	Std. dev.	Min	Max
logrta	2853	11.26	1.36	5.99	17.34
arrps	2851	0.00	0.02	-0.91	0.01
lever	2853	0.60	0.53	-0.19	14.25
current	2853	3.74	21.89	0.00	849.23
cashfps	2851	-0.10	65.37	-2889	1496
solvency	2853	0.53	43.66	-1764	390.20
scps	2851	0.42	8.47	0.00	432.30
vadps	2851	0.01	5.22	-222.60	0.99
insdarr	2853	37.43	9.57	0.28	75.67
indshar	2853	0.00	0.02	0.00	1.00
fshare %	2853	10.26	28.82	0	100
herfind	2853	6189	3531	1	10000

Table 1.4: Descriptive Statistics of the Firm Level Variables, Years 2001-2002

2001

Variable	Obs	Mean	Std. dev.	Min	Max
logrta	1773	11.54	1.42	3.56	16.83
arrps	1772	0.00	0.00	-0.16	0.02
lever	1773	0.53	0.36	-0.05	4.09
current	1773	4.44	28.75	0.00	1123
cashfps	1772	-0.85	37.47	-1567	89.29
solvency	1773	0.57	32.73	-790.35	381.51
scps	1772	0.40	4.61	0.00	129.00
vadps	1772	-0.04	7.40	-296.43	0.99
insdarr	1773	24.83	13.04	0.03	124.75
indshar	1773	0.00	0.02	0.00	0.79
fshare %	1773	11.66	30.57	0	100
herfind	1773	6076	3508	2	10000

2002

Variable	Obs	Mean	Std. dev.	Min	Max
logrta	545	11.82	1.39	7.62	17.37
arrps	543	0.00	0.00	-0.01	0.00
lever	545	0.49	0.44	-0.01	5.59
current	545	4.70	13.54	0.04	191.28
cashfps	543	0.61	19.87	-162.04	432.00
solvency	545	-1.18	31.20	-435.66	108.23
scps	543	1.10	20.63	0.00	481.00
vadps	543	-22.67	532.23	-12402	0.96
insdarr	545	28.23	20.41	0.08	71.95
indshar	545	0.01	0.06	0.00	0.67
fshare %	545	12.42	31.32	0	100
herfind	545	5982	3505	9	10000

Table 1.5: International Investor Sentiment

year	Czech R.	Hungary	Poland	*sentim*
1995	.33	.33	.50	1.17
1996	.42	.42	.42	1.25
1997	.58	.83	.75	2.17
1998	.50	.58	.42	1.50
1999	.58	.58	.67	1.83
2000	.75	.58	.83	2.17
2001	.33	.25	.33	.92
2002	.42	.42	.33	1.17
2003	.67	.50	.50	1.67

Note:
A dummy variable for a given month and country equals 1 if the the monthly market return is higher than the long-term average (1993-2003). *Sentim* is the previous year's average of monthly sums of the three dummies for the Czech Republic, Hungary, and Poland.

Table 1.6: Labor Costs per Value Added in Manufacturing

	1998	2000	2002
Belgium	.62	.61	.63
Czech Republic	.57	.51	.54
Denmark	.69	.68	.68
Germany	.72	.73	.72
Greece	.4	.38	.37
Spain	.64	.65	.68
France	.60	.60	.59
Italy	.55	.55	.56
Luxembourg	.59	.61	.63
Hungary	.52	.55	.54
Netherlands	.59	.59	.63
Austria	.63	.57	.56
Poland	.63	.61	.6
Portugal	.60	.61	.63
Slovakia	.54	.48	.53
Finland	.53	.51	.57
Sweden	.61	.61	.66
United Kingdom	.69	.81	.81

Note:

Based on the median firm in NACE sector D (manufacturing)

Table 1.7: Employee Compensation Rate (Foreign Country/ Czech Republic)

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004
Australia	4.32	3.51	3.84	3.77	3.01	3.73	3.24	2.91	2.57	2.76
Austria	6.02	5.13	5.04	5.28	4.70	4.68	4.40	4.01	3.77	3.60
Belgium	7.17	6.14	6.02	6.25	5.69	5.62	5.29	4.88	4.59	4.43
Canada	4.64	3.94	4.10	4.68	3.71	4.67	4.51	4.04	3.20	3.12
Denmark	6.76	6.14	5.84	6.10	5.55	5.62	5.32	4.83	4.60	4.40
Finland	5.78	4.99	4.87	5.27	4.73	4.77	4.57	4.12	3.90	3.77
France	6.49	5.57	5.40	5.59	5.00	4.94	4.63	4.21	3.96	3.81
Germany	6.23	5.32	5.11	5.31	4.70	4.66	4.31	3.89	3.63	3.43
Greece	2.28	2.14	2.27	2.45	2.30	2.35	2.25	2.14	2.04	2.02
Hungary	1.48	1.14	1.13	1.19	.94	1.08	1.08	1.14	1.18	1.09
Iceland	5.02	4.56	4.62	5.44	5.1	5.99	5.38	4.44	4.56	4.23
Ireland	4.71	4.05	4.10	4.36	4.05	4.13	3.96	3.64	3.41	3.36
Italy	5.12	4.53	4.46	4.55	4.09	4.08	3.82	3.49	3.31	3.18
Japan	8.90	6.95	6.16	6.31	5.78	7.43	6.43	5.13	4.32	3.75
Luxembourg	7.20	6.14	5.97	6.28	5.76	5.88	5.55	5.10	4.74	4.56
Mexico	1.04	.71	.80	1.08	.83	1.08	1.16	1.19	.86	.64
Netherlands	5.00	4.29	4.18	4.47	4.05	4.10	3.90	3.67	3.48	3.32
New Zealand	3.03	2.54	2.74	2.66	2.02	2.33	1.95	1.77	1.78	1.77
Norway	6.75	5.95	5.90	6.47	5.46	6.11	5.86	5.61	5.83	4.91
Poland	.88	.91	.93	1.01	.96	1.00	1.11	1.09	.86	.71
Portugal	2.38	2.19	2.17	2.33	2.13	2.21	2.11	1.95	1.85	1.78
Slovak Republic	.71	.63	.67	.79	.67	.76	.70	.69	.67	.70
Spain	4.17	3.72	3.67	3.87	3.48	3.5	3.32	3.09	2.97	2.91
Sweden	6.33	6.06	6.05	6.32	5.14	5.93	5.45	4.67	4.45	4.36
Switzerland	7.69	7.06	6.13	6.58	5.86	5.87	5.82	5.37	5.05	4.45
Turkey	.73	.73	.81	.82	.77	.76	.84	.49	.42	.47
United Kingdom	5.11	4.17	4.67	5.59	4.87	5.66	5.42	5.1	4.56	4.15
United States	6.07	5.02	5.14	6.27	5.41	6.54	6.60	6.37	5.10	4.13

Note:
Source - OECD

Table 1.8: Statutory Corporate Income Tax Rates (%)

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004
Austria	34	34	34	34	34	34	34	34	34	34
Belgium	40.2	40.2	40.2	40.2	40.2	40.2	40.2	40.2	34	34
Denmark	34	34	34	34	32	32	30	30	30	30
Finland	25	28	28	28	28	29	29	29	29	29
France	36.7	36.7	36.7	41.7	40	36.7	36.4	35.4	35.4	35.4
Germany	56.8	56.7	56.7	56	51.6	51.6	38.3	38.3	39.6	38.3
Greece	40	40	40	40	40	40	37.5	35	35	35
Ireland	40	38	36	32	28	24	20	16	12.5	12.5
Italy	52.2	53.2	53.2	41.3	41.3	41.3	40.3	40.3	38.3	37.3
Luxembourg	40.9	40.9	39.3	37.5	37.5	37.5	37.5	30.4	30.4	30.4
Netherlands	35	35	35	35	35	35	35	34.5	34.5	34.5
Portugal	39.6	39.6	39.6	37.4	37.4	35.2	35.2	33	33	27.5
Spain	35	35	35	35	35	35	35	35	35	35
Sweden	28	28	28	28	28	28	28	28	28	28
United Kingdom	33	33	31	31	30	30	30	30	30	30
Cyprus	25	25	25	25	25	29	28	28	15	15
Czech Republic	41	39	39	35	35	31	31	31	31	28
Estonia	26	26	26	26	26	26	26	26	26	26
Hungary	19.6	19.6	19.6	19.6	19.6	19.6	19.6	19.6	19.6	17.7
Latvia	25	25	25	25	25	25	25	22	19	15
Lithuania	29	29	29	29	29	24	24	15	15	15
Malta	35	35	35	35	35	35	35	35	35	35
Poland	40	40	38	36	34	30	28	28	27	19
Slovakia	40	40	40	40	40	29	29	25	25	19
Slovenia	25	25	25	25	25	25	25	25	25	25

Note:
Source - European Comission

Table 1.9: Im - Pesaran - Shin Tests for Panel Unit Roots

Variable	trend	$t-bar$	Z_{t-bar}	P-value
fshare	no	-3.06	-20.15	0.00
fshare	yes	-0.93	-14.75	1.00
herfind	no	-1.40	1.09	0.86
herfind	yes	-2.53	4.30	0.00
insdarr	no	-2.37	-11.26	0.00
insdarr	yes	-2.27	-1.16	0.12

Note:
$t-bar$ mean of the individual Dickey-Fuller t-statistics of each unit in the panel
$Z_{t-bar} \sim$ N(0,1) under the null hypothesis of non-stationarity

Table 1.10: Hurlin Tests for Homogeneous Non-Causality in Panel Data

H_0	Z_{NT}^{HNC}	P-value
herfind does not Granger cause fshare	-5.27	0.00
fshare does not Granger cause herfind	1.539e+14	0.00
insdarr does not Granger cause fshare	-4.13	0.00
fshare does not Granger cause insdarr	-6.00	0.00

Table 1.11: Specification Tests

H_0	H_A	Hausman's χ^2	P-value	Endogenous Vars
LM-RE	LM-FE	161.88	0.00	N/A
LM-FE	LM-FE-2SLS	24.58	0.02	herfind, insdarr
LM-FD	LM-FD-2SLS	15.08	0.18	herfind, insdarr
LM-FE	LM-FE-AR	345.65	0.00	N/A

Table 1.12: OLS and Fixed Effects Models

	LM-OLS (1)	LM-FE (2)	LM-FE-2SLS (3)
year	.174*	.782***	.892***
sentim	-.865***	.519***	.689***
linta	-6.849*	30.53***	41.854***
ecra	21.072***	18.399***	18.1***
dcita	.272***	.123***	.15***
insdarr	.014**	.03***	.06***
scps	-.013	.009	-.009
vadps	-.0002	-.007	-.019
arrps	1.152	.478	1.571
lever	.218	-.098	-.085
current	.0009	.0004	.0003
cashfps	.0001	-.0005	-.0005
solvency	-.0005	-.002*	-.002*
logrta	.139	.383*	.374*
herfind	.0007***	.0004***	.0008**
indshar	14.033***	23.141***	37.169***
cons	-369.182*	-1604.679***	-1834.043***
rho	.858***	.066***	.034***
Obs.	12460	12460	11138
R^2	.825	.8205	.8084
F-stat.	3665.953***	866.218***	45109.40***

Note: Stars denote significance at 90, 95, and 99%.

Table 1.13: First Differences Models

	LM-FD (1)	LM-FD-2SLS (2)
D.sentim	.321**	.496***
D.linta	16.519***	24.201***
D.ecra	19.322***	19.249***
D.dcita	-.212***	-.179***
D.insdarr	.024***	.038***
D.scps	.022	.036
D.vadps	.016	.023
D.arrps	-1.002	-1.489
D.lever	-.138	-.113
D.current	.0003	.0002
D.cashfps	-.0003	-.0007
D.solvency	-.001	-.001
D.logrta	.054	.075
D.herfind	.0004***	.0001
D.indshar	2.885	4.207
cons	.549***	.673***
Obs.	7261	6424
R^2	.673	.8169
F-stat.	996.165***	13232.23***

Note: Stars denote significance at 90, 95, and 99%.

Table 1.14: Panel Data Logit, Probit, and Tobit Models

	Pro-RE (1)	Lo-FE (2)	To-RE (3)
year	.25***	.003	6.066***
sentim	-.057	-.105	1.665
linta	.367	-2.881	68.98***
ecra	2.884***	1.703***	36.266***
dcita	.027	.197*	-.335***
insdarr	.011**	.019	.13***
scps	-.01	13.739	-.235
vadps	-.00009	1.162	-.006
arrps	.505	270.995	7.389
lever	.006	2.075	.108
current	.00006	.012	.001
cashfps	.0001	1.105	.001
solvency	0	-.016*	-.021**
logrta	.149*	.539	.181
herfind	.0007***	.0005***	.006***
indshar	5.889**	90.348*	92.342***
cons	-517.509***	.	-12326.7***
Obs.	12460	399	12460
Chi^2 (Wald, LR, Wald)	286.94***	241.47***	6602.78***

Note: Stars denote significance at 90, 95, and 99%.

Table 1.15: Cox's Proportional Hazard Model

	PH-Cox (1)
insdarr	.008**
scps	.007
vadps	.002
arrps	-.524
lever	-.111
current	-.012
cashfps	.0008
solvency	-.0004
logrta	.448***
herfind	.0001***
indshar	.923
cons	.
Obs.	13973
Wald Chi^2	259.86***

Note: Stars denote significance at 90, 95, and 99%.

Table 1.16: Truncated Dependent Variable Model for Foreign Stakeholders

	Truncated dependent variable model (1)
sentim	-.432
lintensi	-.617
ecr	-.01
dcit	.013
insdarr	.047***
scps	.061
vadps	-.089
arrps	-29.954
lever	-.987
current	-.00009
cashfps	-.002
solvency	.002
logrta	-.263
herfind	.009***
indshar	-.187
cons	10.286*
Obs.	1877
Wald Chi^2	7518.02***

Note: Stars denote significance at 90, 95, and 99%.

Figure 1.1: Histogram, Accounting Rate of Return

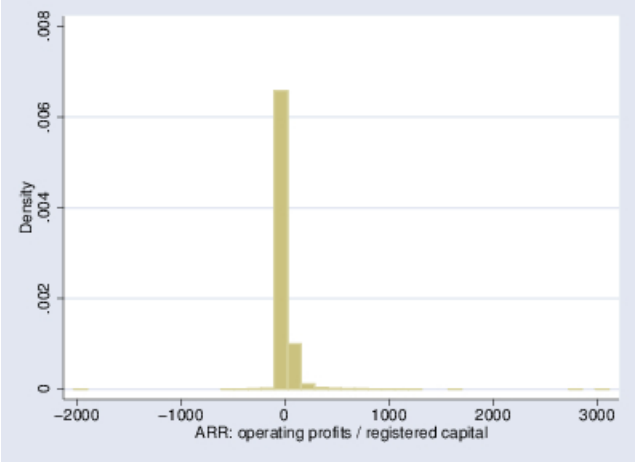


Figure 1.2: Histogram, Ownership by Foreigners, %

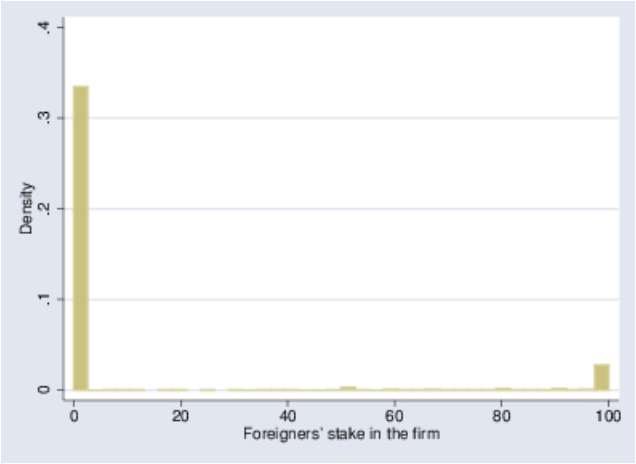
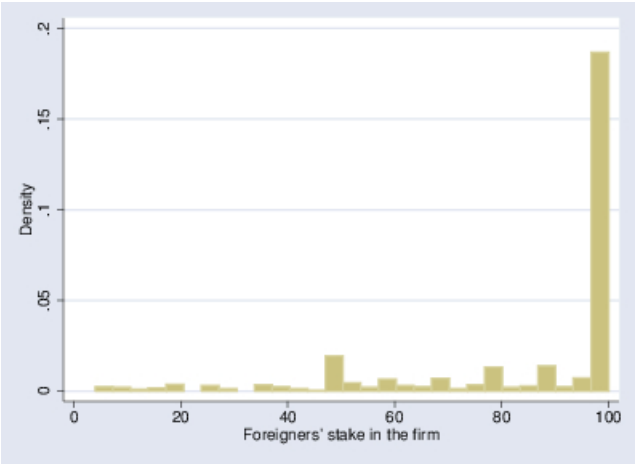


Figure 1.3: Histogram, Ownership by Foreigners, Positive Values, %



Chapter 2

External finance and the introduction of the euro: firm level evidence on debt and equity issues[1]

Abstract

Could eurozone firms raise more external finance thanks to lower transaction costs after the introduction of the euro? Using a panel of about 6,000 Western European listed firms observed from 1995 to 2002, we estimate the likelihood of issuing equity or debt before and after eurozone entry and compared to firms outside the eurozone. At the same time, individual leverage targets, firm size and profitability are taken into account. We find a positive euro effect on the likelihood of issuing debt and external finance in general. A positive effect in case of equity issuance holds only for industries with higher external finance dependence. The additional findings are consistent with previous studies and the trade-off theory of capital structure. Namely, firms tend to revert to their leverage targets, more profitable firms are more likely to issue debt and less likely to issue equity, holding their growth opportunities constant.

JEL classification: G32, F36
Keywords: European single currency, external finance, corporate debt issuance, equity issuance, capital structure, target leverage

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2.1 Introduction

The introduction of the common European currency — the euro — on the 1st of January 1999 was a major event in the international finance. The euro-area financial market has emerged as second largest in the world. The launch of euro was agreed already in 1992 by the Maastricht Treaty. Subsequently, eleven countries adopted euro in 1999, six more have joined during the 2000s. The adoption of the euro seems to be an appealing choice for many countries in the European Union (EU) even though the recent economic crises has tested the currency and shown the pitfalls for countries with loose fiscal policies.

The current study focuses on the importance of the introduction of the euro on firms' external financing decisions. We study corporate debt and equity issuance in the period before and after the launch of the euro, which is from1995 to 2002. The emergence of the eurozone could be considered as an event of financial development, which brings a decrease in the costs of capital. This is because of the abolishment of exchange rate risks and transaction costs related to currency conversion between the euro area member countries. Therefore, it would be natural to believe that many firms with growth opportunities in the eurozone would take advantage of cheaper external finance and issue more debt or equity after 1999. Others would simply adjust to their target leverage ratios by changing the proportions of their debt and equity at a lower transactions cost.

The latter idea of target debt ratios follows capital structure theories. These, such as the trade-off theory or the pecking order theory, look at the costs and benefits of issuing debt and equity to determine their "optimal" proportions. In addition, as Hovakimian, Opler, and Titman (2001) argue, firms' target debt ratios are likely to change over time as their profitabilities and stock prices evolve. At the same time, firms may face impediments to adjust to their target capital structures. We follow the authors and control for deviations from target debt ratios, profitability and the stock market values of firms as potential drivers of debt and equity issues. Therefore we can check previous empirical findings on capital structure choice in the enviroment of introducing the euro.

As regards previous evidence on the "euro effect", some studies document newly emerged opportunitites to diversify investment risks in the eurozone. For example, Hardouvelis, Malliaropulos, and Priestley (2007) show that the home country bias in financial investments in the eurozone has decreased after 1999. Danthine, Giavazzi, and von Thadden (2001) add that pension funds and life insurance companies in many eurozone countries faced certain investment regulations. This required most of their assets to be held in the same currency as their liabilities[2]. Hardouvelis, Malliaropulos, and Priestley (2007) also estimate the effect of launching the euro on the costs of financing. They show that, during the 1990s, the cost of equity financing has dropped by up to 3% for some industries in countries, which adopted euro. A similar decrease in the costs of equity finance was not observed among the European countries which did not join euro. Further, Santos and Tsatsaronis (2003) show that the direct cost of debt financing, i.e. underwriting fees, have been reduced after the euro introduction. The increase in debt financing was also found by Rajan and Zingales (2003), who showed that the ratio of bond issues to GDP has gone up in eurozone countries compared to countries not joining the euro.

The present study complements the existing literature on "the euro effect" by considering firm-level data. Our estimation strategy is focused at external finance issance decisions over time and across firms both in and out of the eurozone. In addition, external finance dependence (EFD) of industries is taken into account. The argument for using EFD is in line with Rajan and Zingales (1998). This means that if the euro introduction has decreased the cost of capital, then firms from industries that rely more on external finance, should be affected disproportionately more in their issuance activities.

We use firm-level data from the Amadeus database provided by Bureau Van Dijk. The analysis focuses on listed firms from 14 Western European countries, ten of which are founding members of the eurozone. We study the time period from 1995 to 2002, which covers a few years before and after the introduction of the euro. Our estimation

[2]For details see Table 3.2. in the mentioned article.

approach follows Hovakimian, Opler, and Titman (2001), where equity and debt issues are explained by the deviation from target leverage, profitability and the market to book ratio representing growth opportunities. We add a dummy variable to capture the euro introduction for the member countries (euro dummy) and an industry-specific proxy for EFD interacted with the euro dummy.

The results suggest that the euro increased the likelihood of issuing external finance in general, which can be attributed mainly to debt issues. The likelihood of issuing equity in the eurozone increased only in industries with high external finance dependence. Further, the findings are consistent with capital structure theories and previous empirical studies of capital structure (Hovakimian, Opler, and Titman (2001)). Namely, we provide additional evidence that firms tend to revert to their dynamic target debt ratios. Finally, our estimates also support a trade-off theory story, i.e. more profitable firms are more likely to issue debt and less likely to issue equity, holding their growth opportunities equal.

This paper is organized as follows: in the next section, the working hypothesis and the empirical estimation methodology are introduced. The data section follows. In section 4 the results are presented. In section 5 couple of alternative specifications are tested for confirming the robustness of the results. Finally, section 6 concludes.

2.2 Hypothesis and testing

The introduction of the euro can be viewed as a natural experiment in financial development. Accordingly, we sort firms into control and treatment groups. The latter is composed of publicly traded companies registered in the ten countries, which were founding members of the euro area[3]. In the former we include listed firms from EU member countries outside the euro-area, such as Denmark and Sweden[4] and from Western European countries that are not EU members, such as Norway and Switzerland.

[3]Luxembourg is not considered due to lack of data.

[4]UK firms are excluded, since they were used for calculating the benchmark EFD. In section 5, the robustness of results is checked by including UK firms in the sample.

According to our hypothesis, external finance issuance is more likely for eurozone firms after 1999. We follow the approach of Hovakimian, Opler, and Titman (2001) for studying equity and debt issues and extend this framework by controlling for the euro effect[5]. The idea of the authors is to estimate whether the issue of new capital or retire/repurchase of existing capital takes firms closer to their target leverage ratios, which is consistent with trade-off theory [6]. We expect the deviation from target leverage to have a significant negative coefficient on the likelihood of debt issuance and a significant positive coefficient on the likelihood of equity issuance. The target leverage is obtained from estimating a static leverage regression, where explanatory variables are profitability (defined as profit/losses of period over total assets), tangibility (defined as tangible fixed assets over total assets), log size of the firm, firm-specific fixed effects and year dummies[7]. Leverage is defined as long-term debt over long-term debt plus shareholders capital. It is constrained to have value between 0 and 1 in our analysis. We estimate the following equation:

$$D_{it} = \alpha + \alpha_i + \alpha_t + \beta EURO_{it} + \gamma X_{it-1} + \varepsilon_{it} \tag{2.1}$$

where i is a firm index and t is a year index (t=1996-2002). D_{it} is a dummy variable representing issuing decision. The dummy is used in four versions to consider seperately the issuance of external finance in general, debt, equity and debt rather than equity. In the first case, the dummy is equal to one if the firm issues external finance (equity, debt or both)[8] and zero otherwise. Second, it is equal to one if the firm issues debt and zero if the firm does not issue any external finance (debt or equity). Third, it is equal to one

[5]Similar empirical methods have been used later in Hovakimian (2004) and Hovakimian, Hovakimian, and Tehranian (2004).

[6]Chen and Zhao (2005) show that the leverage can be mean reverting irrespectively of which capital structure theory firm follows. So the results cannot be viewed as the domination of trade-off theory over other theories of capital strucure.

[7]For a similar way of estimating time-variant leverage targets see for example Rajan and Zingales (1995) and Harris and Raviv (1991).

[8]We follow Hovakimian, Opler, and Titman (2001) in defining the firm to be debt or equity issuing. The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. In section 5 the robustness of results are provided while the cut off level is increased to 10%.

if the firm issues equity and zero if the firm does not issue any external finance. In the fourth case, the dummy is equal to one if the firm issues debt and zero if the firm issues equity. $EURO$ is the binary indicator of the euro, which equals to one for eurozone firms after 1999 (1999 included) and zero otherwise. X_{it-1} represents the vector of firm-specific variables: the leverage deviation from its target, profitability[9] and log of total assets. We also control for firm- and year-specific fixed effects. ε_{it} is the error term. The coefficient of our interest, β, should be positive for external finance issues, debt issues and equity issues and its sign is unclear for the choice between debt and equity issues.

In addition, we introduce a proxy for corporate financing needs based on an industry average indicator computed on the sample of UK firms. Financially constrained firms should be those benefiting from the euro introduction the most and should become debt or equity issuers with a higher probability. The need of external finance is proxied by the firm's industry-specific benchmark external finance level. We follow Rajan and Zingales (1998), who show that some industries are in a greater need of outside capital than others. These are usually industries with a more intensive R & D activity, larger scale plant investment needs and smaller regular turnover, ceteris paribus. In their international study, the authors use the US as a benchmark country for determining EFD. They argue that US firms face the least of financing constraints due to well developed financial markets. Therefore the external financing levels in the US express the natural industry demand for external financing. According to the authors' estimates, the most external finance dependent industries in the US during the 1980's were the manufacturing of pharmaceuticals, plastic products, office machinery and computers and radios. The least dependent industries on the other hand were the manufacturing of tobacco, pottery, and leather.

We use the industry average financial dependence ratio of UK firms as a benchmark due to several reasons. First, the UK did not adopt the euro, hence British firms' financing is not affected directly by the 1999 event. Second, the UK is considered to be an economy with the deepest, most liquid and most developed financial markets in Europe. So, the

[9]We exclude the extreme observations — top 1% and bottom 1% of leverage deviation from its target and profitability are excluded.

eurozone is on the way to catch up with the UK in financial development. In other words we assume that the observed external finance dependence of particular UK industries is "optimal", and is determined by technological differences between industries.

Rajan and Zingales (1998) define the external finance dependence ratio as the difference between capital expenditures and cash flow from operations[10] in the nominator and capital expenditures in the denominator. In other words, EFD gives the percentage of capital expenditures that is not financed by cash flows. The higher the EFD, the more external finance is demanded. Our EFD indicator is very similarto the above. As the category of capital expenditures is not available in the data, the change in fixed assets is used instead[11].

To test whether the euro affects firms in industries with greater external finance dependence disproportionately more than firms in other industries, we include the industry-specific EFD indicator interacted with euro dummy in our regression:

$$D_{it} = \alpha + \alpha_i + \alpha_t + \beta EURO_{it} + \delta EURO_{it} EFD_j + \gamma X_{it-1} + \varepsilon_{it} \tag{2.2}$$

where EFD_j is external finance dependence[12] of industry j corresponding to firm i. We expect δ to be positive for the likelihood of external finance, debt and equity issuance. The sign of δ is unclear for the probability of choosing debt rather than equity.

2.3 Data

Our firm-level dataset comes from the Amadeus[13] database collected by Bureau Van Dijk. The database contains European firms' balance sheets and income statements. We focus on firms from 14 Western European countries over the period 1995-2002. Only firms

[10]Cash flow from operations is defined as the sum of cash flow from operations, decrease in inventories, decrease in receivables and increase in payables.

[11]Capital expenditures in Rajan and Zingales (1998) refer to Compustat item #128, which does not take into account the income from selling the fixed assets.

[12]We sum the nominator and denominator of the EFD index over the years for each firm in UK and calculate the ratio by using the sums (so we avoid the possible extreme values). We take the industry median as a proxy for EFD. We exclude from the analysis industries with less than 3 UK firms present.

[13]The acronym stands for Analyse MAjor Databases from EUropean Sources.

listed on the stock market are included. The financial intermediation sector and which were not established before 1999 and which had left our sample by 1999.

Table 2.1 presents summary statistics of the evolution of debt and equity issuance in the euro area and the rest of Western Europe. First, note that the number of firm years is almost 6 times larger in eurozone sample. Also notice that the sample is unbalanced: we have much more observations after the introduction of the euro in 1999. The frequencies of external finance issues have increased over time. The most striking is the increase in the frequency of equity issues from 7% to 20% for eurozone countries (from 9% to 18% for other countries). The frequency of debt issues[14] has been for eurozone and non-eurozone countries (10%(7%) before and 18%(19%) after 1999.

Table 2.2 contains the summary statistics of the main financial variables. Several interesting stylized facts appear. The median eurozone firm is smaller compared to non-eurozone firms both before and after the introduction of the euro. Debt issuing firms are larger, have more tangible assets, are more levered and have a lower market-to-book (MTB) ratio[15] than equity issuing firms. These differences among debt and equity issuers is consistent with evidence in previous studies Hovakimian, Opler, and Titman (2001).

2.4 Results

Table 2.3 presents the OLS estimates of equation (1). In the first column, the dependent variable is the dummy variable equal to 1 if firm issued external finance (debt, equity or both). The introduction of the euro has a positive effect on external finance issues. The firm belonging to the euro-area has a 6% higher probability to issue external finance compared to other firms. The coefficient on the deviation of leverage from target leverage is negative and is statistically significant. This means, if a firm is over-levered then it is less likely to issue external finance in general. We also find that more profitable and

[14]Keep in mind that throughout the paper we refer to debt issues without distinguishing between their public (e.g. bond issue) and private (e.g. bank credit) alternatives.

[15]MTB ratio is defined as book value of liabilities plus market value of equity over total book value of assets.

smaller firms are more likely to issue external finance in general.

Next, we looked at debt and equity issuance separately. Results for debt issues versus neither equity, nor debt issues are reported in the second column of Table 2.3. The introduction of the euro has a positive effect on debt issues. The firm belonging to euro-area has 7% higher probability of issuing debt compared to other firms. The coefficient on the deviation from target leverage is negative and is statistically significant. This implies that over-levered firms are less likely to issue debt and vice-versa. This result confirms a sort of a mean-reverting behavior of leverage. We also find that higher profitability and smaller firm size increase the likelihood of issuing debt.

The results for equity issues versus neither equity, nor debt issues are reported in the third column of Table 2.3. The effect of the euro on equity issues is not statistically significant. The coefficient on deviation from target leverage is positive and statistically significant. This implies that over-levered firms are more likely to issue equity and vice-versa, which supports the results on debt issuance. A significant coefficient is also found on size, meaning that smaller firms are more likely to issue equity.

The last column of Table 2.3 is of interest for disentangling the effect of the introduction of the euro on the choice between issuing debt or equity. We find at 10% significance level that eurozone firms are more likely to issue debt rather than equity after 1999. The coefficient on the deviation from target leverage is again consistent with the mean-reverting behavior of leverage. To sum up the findings above, we can conclude that external finance issuance in the eurozone after 1999 was realized mostly through the debt channel.

Next we estimate equation (2) to see whether the industry specific dependence on external finance matters for the issuing decisions. Table 2.4 presents the OLS estimates of equation (2). Results for the probability of external finance issuance are presented in the first column. Similarly to Table 2.3, the euro has a positive effect on external finance issuance. Moreover, the magnitude of the effect is robust to the inclusion of the interaction term of the euro dummy and the industry-specific EFD indicator. However,

the interaction term has a statistically insignificant coefficient. This is to say that firms from high EFD industries behave the same as firms in low EFD industries, as regards external financing issuance in general.

Results for debt issues versus neither equity, nor debt issues are reported in the second column of Table 2.4. The euro dummy has a positive and significant coefficient, as in Table 2.3, and the euro and EFD interaction term has a negative but insignificant coefficient. Hence all firms, regardless of the industry they belong to, have an increased debt issuance.

Column (3) of Table 2.4 presents the results for equity issues versus neither equity, nor debt issues. Here the euro dummy is still statistically insignificant, as in Table 2.3, but the interaction term of euro dummy and the EFD indicator is positive and statistically significant at the 1% level. Thus the above implies that eurozone firms from industries with higher EFD are more likely to issue equity after 1999.

The last column of Table 2.4 reports the results of debt versus equity issues. The euro dummy has a significantly positive coefficient. That is, the firms are more likely to issue debt than to issue equity in the eurozone after 1999. The interaction term of euro dummy and the EFD indicator has a significantly negative coefficient. This implies that eurozone firms from industries with higher EFD prefered to issue equity instead of debt after the introduction of the euro, as opposed to lower EFD firms[16]. The above result confirms the findings of previous empirical studies on capital structure (Hovakimian, Opler, and Titman (2001)) and is also consistent with the pecking order theory[17]. According to the pecking order theory, internal and external funds have different costs due to the asymmetry of information between the insiders and outsiders of the firm. Therefore, firms will first use the internal funds and among the external funds the "safe" debt is preferred to "risky" equity. In other words — equity would be issued only after the

[16]For calculating the precise effect we need to add up the coefficients in front of the euro dummy and the interaction term. The EFD measure varies from -4.45 for firms in manufacturing of textiles (NACE 2 industry code 17) to 4.07 for firms in research and development (NACE 2 industry code 73). That is, firms from the low EFD textiles industry are 31% (0.115+(-0.044*-4.45)=0.311) more likely to issue debt than equity. On the contrary, firms from the high EFD R & D sector are 6.4% (0.115+(-0.044*4.07)=-0.064) more likely to issue equity rather than debt.

[17]See for example Shyam-Sunder and Myers (1999) and Myers (1984).

exhaustion of internal sources and debt. In our setup, firms from industries with higher EFD are likely to be the ones who have exhausted the debt finance available and therefore turn to the remaining resource of financing — equity.

2.5 Robustness checks

In this section, a couple of alternative specifications are estimated to confirm the robustness of results. First, it might be the case that the change in fixed assets does not measure investments closely enough. We present an alternative EFD measure using tangible fixed assets. In other words, the second EFD ratio (EFD_2) is defined as the change in tangible fixed assets, minus cash flow, plus change in inventories, plus change in receivables, minus change in payables in the nominator and the change in tangible fixed assets in the denominator. The two EFD measures used are highly correlated (71%) with each other. The results with EFD_2 are presented in Table 2.5. The estimates are similar to the ones in Table 2.4. The euro has a positive effect on the probabilities of issuing external finance and debt. The interaction terms are statistically significant in case of equity and the choice between debt and equity issuance, as before.

We also check whether including UK companies in the sample changes the results. As a result, the number of observations increases by roughly 45%. The estimates with UK firms included in the sample are presented in Table 2.6. Again, the results are similar to the ones reported in Table 2.4, except for the coefficients of the euro dummies. Those estimates have diminished slightly according to the first two columns — external finance and debt issues. Further, the euro dummy is statistically insignificant in the last two columns. The interaction terms in the last two columns are statistically significant and of the similar magnitude as in Table 2.4. We can conclude that including UK firms in the sample does not change the results substantially.

There might be a concern that the cut-off level of 5% in defining the firms being debt and equity issuers was chosen arbitrarily. We repeat the estimation with redefining these

dummies and lifting the cut-off value to 10%. To reiterate, the debt issue dummy is now equal to one if the change in the firm's long-term debt to total assets ratio is larger than 10% and zero otherwise. The equity issue dummy is equal to one if the change in the firm's book share capital to total assets is larger than 10% and zero otherwise[18].

The results with the modified dependent variables are reported in Table 2.7. The first three columns show comparable results to Table 2.4. In the final column, the interaction term is not statistically significant anymore, but still has a negative sign. This indicates that firms from high EFD industries behave in a similar manner to firms in other industries after euro introduction and there is no evidence for a preference of acquiring equity finances rather than debt finances.

Finally, it could be argued that it is important to control for growth opportunities in evaluating the firm's external finance decisions. So in Table 2.8, we add an additional control variable, the market-to-book (MTB) ratio. Note that this variable is observed only for two-thirds of the firms, so the sample is reduced quite a bit. The coefficient on MTB has a very low value in all columns and is statistically significant only in column (2). Hence, firms with a higher MTB ratio are more likely to issue debt. Additionally we may note that the coefficient estimates of the euro dummies have increased compared to Table 2.4 in the first two columns. The significance of the interaction terms in the last two columns has decreased, but the signs of the coefficients are in line with the previous results.

Overall the different robustness checks confirm the findings presented in the previous section.

[18]Note that the above sensitivity check also treats potential international differences in accounting and reporting standards. Those may affect the values of total assets through differences in the evaluation methods applied to certain asset classes. The values of book share capital and long-term debt, however, should be comparable internationally. Therefore the censoring value as a percentage of total assets could in principle matter and we may potentially classify "small" issues of external finances as a positive response in some countries, but not in others.

2.6 Conclusion

This paper investigates the effect of the European currency unification on the reduction of external financing costs and the resulting improved ability of firms to issue debt and equity. Our results suggest that, first, eurozone firms are more likely to issue debt and external finance in general after the introduction of the euro in 1999. This implies that the costs of acquiring external finance, especially those of debt, may have diminished due to the introduction of the euro. Second, if the firm belongs to an industry with high external finance dependence, the probability of issuing equity rather than debt increases with the euro effect. This is consistent with the pecking order theory of capital structure. The theory assings greater asymmetries of information to equity finance compared to debt finance. Therefore, equity finance is used only if the financing sources from internal funds and debt have already been exhausted. This could be the case in high EFD industries, where the funding required by new investment projects is typically larger than in other industries, ceteris paribus. The above provides an explanation why the euro effect increases the probability of equity issuance only in the finance-intensive industries.

Apart from the euro effect and the costs of external finance, we also tested some commonly considered factors driving debt and equity issuance. These factors are interesting from the point of view of capital structure theories. We provide evidence from Western Europe that target debt ratios may change over time and firms tend to gravitate to those target values. Consistently with the predictions of the trade-off theory, our estimates also imply that more profitable firms prefer to issue debt rather than equity, keeping their growth opportunities equal. The above findings are in line with previous empirical studies on U.S. data, such as Hovakimian, Opler, and Titman (2001).

TABLE 2.1—NUMBER OF DEBT AND EQUITY ISSUES

	Euro countries					Non-euro countries				
	Equity issues		Debt issues			Equity issues		Debt issues		
	Obs.	Freq.	Obs.	Freq.	Total	Obs.	Freq.	Obs.	Freq.	Total
1996	91	0.03	181	0.07	272	13	0.05	16	0.03	29
1997	201	0.07	222	0.09	423	7	0.02	12	0.02	19
1998	302	0.11	294	0.12	596	55	0.19	90	0.17	145
1999	369	0.13	412	0.17	781	42	0.15	89	0.17	131
2000	565	0.21	425	0.17	990	56	0.20	112	0.21	168
2001	281	0.10	434	0.18	715	28	0.10	96	0.18	124
2002	938	0.34	474	0.19	1412	82	0.29	114	0.22	196
Total/Average	2747	0.14	2442	0.14	5189	283	0.14	529	0.14	812

NOTES:The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. Euro countries are Austria, Belgium, Finland, France, Germany, Ireland, Italy, Netherland, Portugal and Spain. Non-euro countries are Denmark, Norway, Sweden and Switzerland.

2.7 Appendix: Descriptive Statistics and Results Tables

TABLE 2.2—SUMMARY STATISTICS

	Euro countries				Non-euro countries			
Before 1999								
Equity issuers								
	Mean	Median	St. dev.	Obs.	Mean	Median	St. dev.	Obs.
Total assets	315	28	1488	594	376	37	1420	75
Leverage	0.11	0.00	0.18	561	0.21	0.16	0.24	49
Lev-TargetLev	-0.03	-0.08	0.16	549	0.04	-0.02	0.20	48
ROA	0.08	0.06	0.13	581	-0.02	0.05	0.28	74
MTB	3.05	1.75	7.20	240	2.42	1.47	2.31	40
Tangibility	0.22	0.16	0.22	592	0.19	0.12	0.23	75
Debt issuers								
	Mean	Median	St. dev.	Obs.	Mean	Median	St. dev.	Obs.
Total assets	1396	133	5930	697	1330	214	4311	118
Leverage	0.40	0.36	0.27	697	0.41	0.41	0.21	118
Lev-TargetLev	0.20	0.17	0.26	688	0.21	0.20	0.20	117
ROA	0.04	0.04	0.10	691	0.02	0.05	0.16	117
MTB	1.78	1.37	1.45	364	2.12	1.33	3.68	80
Tangibility	0.36	0.33	0.24	694	0.36	0.29	0.27	118
No issuers								
	Mean	Median	St. dev.	Obs.	Mean	Median	St. dev.	Obs.
Total assets	1239	54	10100	6550	1079	127	3962	1426
Leverage	0.13	0.00	0.29	6218	0.27	0.22	0.30	1022
Lev-TargetLev	-0.04	-0.09	0.31	6168	0.07	0.01	0.28	1015
ROA	0.04	0.05	1.27	6530	0.03	0.04	0.20	1423
MTB	3.24	1.26	33.75	2174	2.27	1.27	4.87	712
Tangibility	0.24	0.20	0.22	6507	0.31	0.26	0.26	1386
After 1999								
Equity issuers								
	Mean	Median	St. dev.	Obs.	Mean	Median	St. dev.	Obs.
Total assets	180	15	1829	2153	315	37	1595	208
Leverage	0.07	0.00	0.34	2017	0.16	0.02	0.23	179
Lev-TargetLev	-0.08	-0.11	0.33	2002	-0.03	-0.11	0.21	177
ROA	-0.06	0.00	0.41	2134	-0.21	-0.01	0.49	208
MTB	3.66	1.73	11.47	1001	2.04	1.28	2.35	151
Tangibility	0.11	0.02	0.18	2148	0.19	0.07	0.26	206
Debt issuers								
	Mean	Median	St. dev.	Obs.	Mean	Median	St. dev.	Obs.
Total assets	2703	174	12300	1745	1192	201	3845	411
Leverage	0.40	0.40	0.43	1745	0.41	0.40	0.19	411
Lev-TargetLev	0.19	0.18	0.43	1744	0.19	0.18	0.18	409
ROA	-0.01	0.03	0.40	1744	-0.01	0.03	0.27	411
MTB	1.85	1.30	1.81	1541	2.11	1.22	3.20	355
Tangibility	0.29	0.24	0.23	1745	0.36	0.31	0.28	409
No issuers								
	Mean	Median	St. dev.	Obs.	Mean	Median	St. dev.	Obs.
Total assets	1065	38	6487	10456	992	103	3875	2325
Leverage	1.80	0.00	168.51	9876	0.17	0.12	1.05	1828
Lev-TargetLev	1.64	-0.10	168.75	9847	-0.03	-0.07	1.06	1821
ROA	0.00	0.02	0.21	10432	-0.05	0.03	0.37	2322
MTB	2.65	1.33	16.68	6451	2.21	1.25	4.12	1888
Tangibility	0.17	0.09	0.21	10434	0.25	0.18	0.24	2316

NOTES: The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. Euro countries are Austria, Belgium, Finland, France, Germany, Ireland, Italy, Netherlands, Portugal, and Spain. Non-euro countries are Denmark, Norway, Sweden and Switzerland. Total assets are given in millions of 1995 US dollars. Leverage is defined as long-term debt over long-term debt plus book share capital. Target leverage is a fitted value from leverage regression. ROA is defined as profit/losses of period over total assets. MTB ratio is defined as book value of liabilities plus market value of equity over total book value of assets. Tangibility is defined as a ratio of tangible fixed assets over total assets.

Table 2.3 — Issue of external finance
Controlling for Euro dummy

	Issue vs.No	Debt vs.No	Equity vs.No	Debt vs.Equity
Euro	.063	.071	.008	.124
	(.022)***	(.021)***	(.014)	(.069)*
Leverage-Leverage target	-.64	-.807	.094	-.552
	(.034)***	(.034)***	(.026)***	(.085)***
ROA	.243	.244	.041	-.008
	(.05)***	(.043)***	(.044)	(.18)
Log(Total assets)	-.167	-.049	-.149	.111
	(.01)***	(.011)***	(.01)***	(.033)***
Const.	2.235	.747	1.855	-.549
	(.118)***	(.134)***	(.118)***	(.374)
Obs.	13344	12260	10971	2901
R^2	.082	.087	.075	.116

Notes: The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. The firm is external finance issuing if firm is either or both debt and equity issuing. Euro dummy equals one for eurozone firms after 1999 (1999 included) and zero otherwise. Leverage is defined as long-term debt over long-term debt plus book share capital. Target leverage is a fitted value from leverage regression. ROA is defined as profit/losses of period over total assets. (Leverage - target leverage), ROA and Log(Total assets) are one period lagged. Robust standard errors are in brackets. ***, ** and * denotes significance at the 1, 5 and 10 percent level respectively. All regressions include firm and year fixed effects.

TABLE 2.4 — ISSUE OF EXTERNAL FINANCE, CONTROLLING FOR EURO DUMMY AND EFD

	Issue vs.No	Debt vs.No	Equity vs.No	Debt vs.Equity
Euro	.065	.069	.013	.115
	(.022)***	(.021)***	(.014)	(.07)*
Euro*EFD	.009	-.004	.018	-.044
	(.007)	(.007)	(.005)***	(.022)**
Leverage-Leverage target	-.638	-.805	.094	-.556
	(.034)***	(.034)***	(.026)***	(.085)***
ROA	.243	.245	.041	-.007
	(.05)***	(.043)***	(.044)	(.179)
Log(Total assets)	-.169	-.048	-.151	.118
	(.01)***	(.011)***	(.01)***	(.033)***
Const.	2.25	.737	1.886	-.614
	(.118)***	(.134)***	(.118)***	(.373)*
Obs.	13328	12244	10958	2898
R^2	.082	.087	.077	.12

NOTES: The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. The firm is external finance issuing if firm is either or both debt and equity issuing. Euro dummy equals one for eurozone firms after 1999 (1999 included) and zero otherwise. EFD is UK industry median EFD defined as (change in fixed assets - cash flow + change in inventories + change in receivables - change in payables)/ absolute value of change in fixed assets. Leverage is defined as long-term debt over long-term debt plus book share capital. Target leverage is a fitted value from leverage regression. ROA is defined as profit/losses of period over total assets. (Leverage - Leverage target), ROA and Log(Total assets) are one period lagged. Robust standard errors are in brackets. ***, ** and * denotes significance at the 1, 5 and 10 percent level respectively. All regressions include firm and year fixed effects.

TABLE 2.5 — ISSUE OF EXTERNAL FINANCE, CONTROLLING FOR EURO DUMMY AND EFD_2

	Issue vs.No	Debt vs.No	Equity vs.No	Debt vs.Equity
Euro	.059	.074	.001	.149
	(.023)***	(.022)***	(.014)	(.07)**
Euro*EFD_2	.001	0	.003	-.008
	(.002)	(.002)	(.0009)***	(.005)*
Leverage - Leverage target	-.638	-.804	.093	-.551
	(.034)***	(.034)***	(.026)***	(.085)***
ROA	.244	.244	.043	-.014
	(.05)***	(.043)***	(.044)	(.179)
Log(Total assets)	-.168	-.048	-.15	.118
	(.01)***	(.011)***	(.01)***	(.033)***
Const.	2.242	.737	1.87	-.623
	(.118)***	(.134)***	(.119)***	(.371)*
Obs.	13328	12244	10958	2898
R^2	.082	.087	.076	.12

NOTES: The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. The firm is external finance issuing if firm is either or both debt and equity issuing. Euro dummy equals one for eurozone firms after 1999 (1999 included) and zero otherwise. EFD_2 is UK industry median EFD defined as (change in tangible fixed assets - cash flow + change in inventories + change in receivables - change in payables)/ absolute value of change in tangible fixed assets. Leverage is defined as long-term debt over long-term debt plus book share capital. Target leverage is a fitted value from leverage regression. ROA is defined as profit/losses of period over total assets. (Leverage - Leverage target), ROA and Log(Total assets) are one period lagged. Robust standard errors are in brackets. ***, ** and * denotes significance at the 1, 5 and 10 percent level respectively. All regressions include firm and year fixed effects.

TABLE 2.6 — ISSUE OF EXTERNAL FINANCE,
CONTROLLING FOR EURO DUMMY, EFD AND UK FIRMS INCLUDED

	Issue vs.No	Debt vs.No	Equity vs.No	Debt vs.Equity
Euro	.04	.041	.001	.018
	(.014)***	(.013)***	(.009)	(.039)
Euro*EFD	.008	-.004	.017	-.042
	(.007)	(.007)	(.005)***	(.022)**
Leverage - Leverage target	-.622	-.766	.082	-.516
	(.029)***	(.029)***	(.023)***	(.066)***
ROA	.185	.259	-.048	.088
	(.038)***	(.033)***	(.033)	(.11)
Log(Total assets)	-.164	-.057	-.13	.115
	(.008)***	(.008)***	(.008)***	(.024)***
Const.	2.113	.816	1.631	-.617
	(.091)***	(.1)***	(.089)***	(.27)**
Obs.	19212	17690	15604	4306
R^2	.076	.076	.067	.118

NOTES: The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. The firm is external finance issuing if firm is either or both debt and equity issuing. Euro dummy equals one for eurozone firms after 1999 (1999 included) and zero otherwise. EFD is UK industry median EFD defined as (change in fixed assets - cash flow + change in inventories + change in receivables - change in payables)/ absolute value of change in fixed assets. Leverage is defined as long-term debt over long-term debt plus book share capital. Target leverage is a fitted value from leverage regression. ROA is defined as profit/losses of period over total assets. (Leverage - Leverage target), ROA and Log(Total assets) are one period lagged. Robust standard errors are in brackets. ***, ** and * denotes significance at the 1, 5 and 10 percent level respectively. All regressions include firm and year fixed effects.

TABLE 2.7 — ISSUE OF EXTERNAL FINANCE,
CONTROLLING FOR EURO DUMMY AND EFD,
INCREASING THE ISSUE DEFINITION CUT OFF TO 10%

	Issue vs.No	Debt vs.No	Equity vs.No	Debt vs.Equity
Euro	.065	.077	-.002	.191
	(.019)***	(.018)***	(.009)	(.095)**
Euro*EFD	.002	-.004	.011	-.028
	(.006)	(.005)	(.003)***	(.032)
Leverage - Leverage target	-.559	-.687	.085	-.513
	(.031)***	(.031)***	(.021)***	(.131)***
ROA	.157	.145	.03	-.064
	(.044)***	(.037)***	(.038)	(.317)
Log(Total assets)	-.157	-.052	-.12	.111
	(.009)***	(.01)***	(.009)***	(.046)**
Const.	1.99	.697	1.465	-.447
	(.109)***	(.116)***	(.108)***	(.502)
Obs.	13328	12244	10958	1905
R^2	.089	.092	.075	.144

NOTES: The firm is debt issuing in given year if the change in long-term debt is larger than 10% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 10% of total assets. The firm is external finance issuing if firm is either or both debt and equity issuing. Euro dummy equals one for eurozone firms after 1999 (1999 included) and zero otherwise. EFD is UK industry median EFD defined as (change in fixed assets - cash flow + change in inventories + change in receivables - change in payables)/ absolute value of change in fixed assets. Leverage is defined as long-term debt over long-term debt plus book share capital. Target leverage is a fitted value from leverage regression. ROA is defined as profit/losses of period over total assets. (Leverage - Leverage target), ROA and Log(Total assets) are one period lagged. Robust standard errors are in brackets. ***, ** and * denotes significance at the 1, 5 and 10 percent level respectively. All regressions include firm and year fixed effects.

Table 2.8 — Issue of external finance,
Controlling for euro dummy, EFD and MTB

	Issue vs.No	Debt vs.No	Equity vs.No	Debt vs.Equity
Euro	.088	.094	.015	.167
	(.035)**	(.035)***	(.018)	(.089)*
Euro*EFD	-.004	-.016	.011	-.01
	(.011)	(.011)	(.006)*	(.033)
Leverage - Leverage target	-1.068	-1.144	.026	-.503
	(.049)***	(.05)***	(.029)	(.127)***
ROA	.13	.25	-.11	.107
	(.068)*	(.06)***	(.057)**	(.318)
Log(Total assets)	-.146	-.028	-.129	.111
	(.018)***	(.02)	(.015)***	(.051)**
MTB	.0003	.001	.0005	-.0002
	(.0004)	(.0002)***	(.0007)	(.008)
Const.	1.965	.503	1.607	-.237
	(.215)***	(.238)**	(.177)***	(.637)
Obs.	8726	8252	7099	1833
R^2	.126	.131	.06	.086

Notes: The firm is debt issuing in given year if the change in long-term debt is larger than 5% of total assets. The firm is equity issuing in given year if the change in book share capital is larger than 5% of total assets. The firm is external finance issuing if firm is either or both debt and equity issuing. Euro dummy equals one for eurozone firms after 1999 (1999 included) and zero otherwise. EFD is UK industry median EFD defined as (change in fixed assets - cash flow + change in inventories + change in receivables - change in payables)/ absolute value of change in fixed assets. Leverage is defined as long-term debt over long-term debt plus book share capital. Target leverage is a fitted value from leverage regression. ROA is defined as profit/losses of period over total assets. MTB ratio is defined as book value of liabilities plus market value of equity over total book value of assets. (Leverage - Leverage target), ROA, Log(Total assets) and MTB are one period lagged. Robust standard errors are in brackets. ***, ** and * denotes significance at the 1, 5 and 10 percent level respectively. All regressions include firm and year fixed effects.

Chapter 3

Currency Shocks to Export Sales of Importers: A Heterogeneous Firms Model and Czech Micro Estimates[1]

Abstract

To what extent can exporters cushion the impact of currency appreciation shocks by using imported intermediates? We apply a partial equilibrium model with heterogeneous firms. Producers can serve the domestic market, export final goods, or import inputs. In the model, an exogenous exchange rate shock simultaneously affects the variable costs and revenues associated with exports and imports. The impact of a hypothetical 1% appreciation of the domestic currency on sales is estimated using a panel of 7,356 Czech manufacturing firms observed from 2003 to 2006. We focus on the above period to exploit the rich within-firm variation in trade strategies. The variation is probably associated with the lifting of trade barriers due to Czech EU membership since 2004. For firms that both export and import, the model predicts a drop in total sales of 0.2%, a drop in export sales of 0.8%, and a rise in domestic sales of 0.2%.

JEL classification: C23, C26, D22, D24, F12.
Keywords: Exchange rate pass-through, heterogeneous firms, international trade, monopolistic competition, production function, total factor productivity

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3.1 Introduction

Over recent years in the Czech Republic, we have witnessed anecdotal evidence of domestic currency appreciation bubbles causing alarm among heads of large export-oriented industrial companies and industrial associations. These managers argued that a strong domestic currency wiped out 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. In this paper we ask to what extent do cheaper imported intermediate products compensate for a drop in export sales as a result of an 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 the 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 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 to engage in all of the mentioned 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 probably 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 3.2 reviews the relevant literature, Section 3.3 sets up the model, Section 3.4 outlines the testable implications of the model, Section 3.5

describes the dataset, Section 3.6 explains the estimation procedure, Section 3.7 summarizes the results, and the last section concludes.

3.2 Literature Review

Based on theory and empirical evidence, more productive and larger firms are more likely to import and export 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, the authors address the simultaneity issue of inputs and the productivity shock by 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 just 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.

[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.

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.

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, the model of Bas and Strauss-Kahn (2011) 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. They do not estimate structural parameters, either.

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, 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.

3.3 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 also decide to use domestic intermediate goods only (M=0) or to employ a mix of domestic and imported intermediates ($M=1$). Firms' export and import decisions 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:

$$c(X=0, M=0) = c, \qquad f(X=0, M=0) = f,$$

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

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

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

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

[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:

$H = \sum^{N}_{i=1}(s^2_i)$, 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 $\boldsymbol{x}$ within a sector takes the standard form:

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:

$$q_{ij} = A_j p_{ij}^{-\varepsilon} \qquad (3.1)$$

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:

$$q_i = e_i(M) I_i \qquad (3.2)$$

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 (3.1), production (3.2), and cost functions $c(X,M)$ and $f(X,M)$ we can write firm i's profit from serving market j as:

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

$$= A_j p_{ij}^{1-\varepsilon} - c(X,M) A(X) p_{ij}^{-\varepsilon}/e_i(M) - f(X,M) \qquad (3.3)$$

The profit-maximizing unit price then becomes:

$$p_{ij}^* = p_i^* = \varepsilon c(X,M)/[e_i(M)(\varepsilon-1)] \qquad (3.4)$$

Plugging the above equilibrium prices (3.4) into the profit function (3.3) we get the following equilibrium profits for various trade strategies:[7]

$$\Pi_i^*(X,M) = \Pi_{i0}^*(M) + \Pi_{ix}^*(M) \qquad (3.5)$$

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

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

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

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

$u(\mathbf{x}) = (x_1^{\alpha} + x_2^{\alpha} + \ldots + x_h^{\alpha})^{1/\alpha}$, 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 3.1 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.

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

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 3.1). Notably, we find firms in our data selecting into all four *(X,M)* strategies within each manufacturing subsector.[8] So we focus on a set of parameters that implies the existence of all trade strategies in sectoral equilibrium.

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 what follows 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.

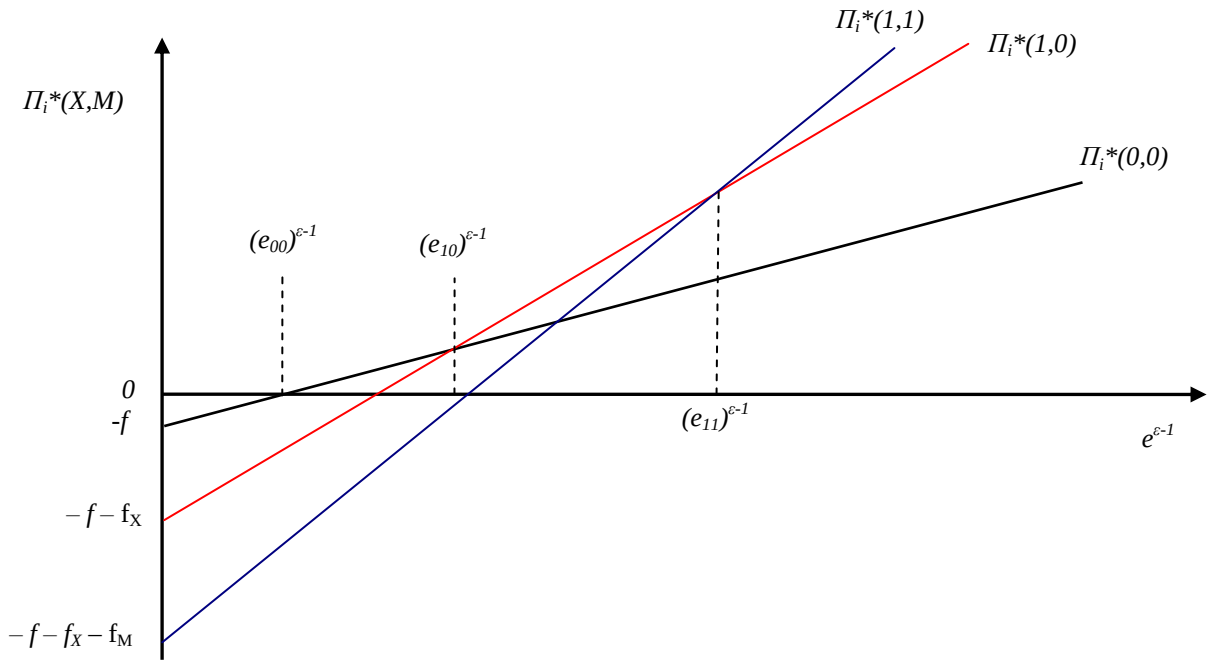
If $\Pi_i^*(0,0)$ is to earn positive profits, productivity e_i must exceed the cut-off point $(e_{00})^{\varepsilon-1} = (fc^{\varepsilon-1}) / EA$. Given that $\Pi_i^*(0,1)$ and $\Pi_i^*(1,0)$ have a lower intercept than $\Pi_i^*(0,0)$, strategies *(0,1)* and *(1,0)* will exist in equilibrium only 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 3.1 depicts the sectoral equilibrium with profit lines for different trade strategies.

[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.

[9] See sales, real value added, real capital, labor, energy, and material inputs in Table 4 in the Data section and Table 3.A1 in Appendix 1.

Figure 3.1: 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 $f_X = f_M$.

3.4 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 (3.1) and (3.4), the equilibrium sales equation of firm i serving market j can be written as:

$$S_{ij}(X,M) = A_j(p_{ij}^*)^{1-\varepsilon} = A_j E' c(X,M)^{1-\varepsilon} e_i(M)^{\varepsilon-1} \qquad (3.6)$$

where $E' = [\varepsilon/(\varepsilon-1)]^{1-\varepsilon}$ is a positive constant. Using (3.6) we can also write total sales in all markets served as a function of trade strategies:

$$S_i(X,M) = S_{i0}(X,M) + S_{ix}(X,M) \qquad (3.7)$$

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

[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.

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

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

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

Now let us introduce the exchange rate into the above sales equations with the aim to estimate 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 (3.6) and (3.7). 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 (3.7) as follows:

$$S_i(0,1) = S_{i0}(0,1) = AE'[c\tau_M r]^{1-\varepsilon} e_i(1)^{\varepsilon-1} \qquad (3.8)$$

[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.

$$S_i(1,0) = S_{i0}(1,0) + S_{ix}(1,0) = (A+rA_x\tau_X^{1-\varepsilon})E'c^{1-\varepsilon}e_i(0)^{\varepsilon-1} \quad (3.9)$$

$$S_i(1,1) = S_{i0}(1,1) + S_{ix}(1,1) = (A+rA_x\tau_X^{1-\varepsilon})E'[c\tau_M r]^{1-\varepsilon}e_i(1)^{\varepsilon-1} \quad (3.10)$$

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:

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

$$\rho_0(0,1) = \rho(0,1) = \rho_0(1,1) = (1-\varepsilon) \quad (3.11)$$

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

$$\rho_x(1,1) = (2-\varepsilon) \quad (3.12)$$

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

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

$$= 1 - \varepsilon + R \quad (3.13)$$

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:

$$\varepsilon^* > 1 + R \quad (3.14)$$

As will be shown, the above condition (3.14), parameter ε, and the listed partial effects (3.11)–(3.13) can be estimated from our data on Czech manufacturing firms. So, finally, we will test the hypothesis that the terms (3.11)–(3.13) are significantly different from zero.

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

To proceed, we take natural logarithms from the equilibrium sales equations (3.7)–(3.10) 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:

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

$$+ (\varepsilon-1)log(e_i(M)) \qquad (3.15)$$

In order to convert (3.15) into an estimable format, let us assume that all the addends in (3.15) 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 (3.15) can be rewritten as follows:

$$s_{it} = \alpha_0 + \alpha_1 d(1,.)_{it} + \alpha_2 d(.,1)_{it} + \alpha_3 TFP_{it} + \theta_{it} \qquad (3.16)$$

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 (3.15), 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 (3.15) are stacked into constants α_0 to α_3 of (3.16) as shown by the following expressions:

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

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

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

$$\alpha_3 = \varepsilon-1$$

which leads to:

$$\varepsilon = \alpha_3 + 1$$

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

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

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

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

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

[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.

Furthermore, based on (3.11), (3.12), and (3.13), 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 (3.16):

$$\rho_0(0,1) = \rho(0,1) = \rho_0(1,1) = -\alpha_3 \qquad (3.17)$$

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

$$\rho_x(1,1) = 1 - \alpha_3 \qquad (3.18)$$

$$\rho(1,1) = 1 - \alpha_3 - exp(-\alpha_1) \qquad (3.19)$$

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.

3.5 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. Looking at Table 1 we see a 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 3.A1 and Table 3.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

[16] See the sales equation (3.16).

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 3.2, 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 3.2, up to 47% of trade strategy shifts lead to strategy *d(1,1)* of both exporting and importing. At the same time, Table 3.3 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]

One of the key building blocks of the model in Section 3.3 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 3.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.

$$logX_{it} = a_0 + a_1 d(0,1)_{it} + a_2 d(1,0)_{it} + a_3 d(1,1)_{it} + A_4 Z_{it} + \omega_{it}$$

The estimates of the above equation can be found in Table 3.4. 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 3.A1 in Appendix 1 we find consistency with the model's assumptions in most cases.[18]

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

[18] The purpose of the exercise was merely to describe the data and perform a consistency check of the model's assumptions. Therefore, the estimates in Table 3.4 should be interpreted as stylized facts without the ambition to test causal relationships. In the latter case we would have had to specify other firm characteristics as explanatory variables.

Table 3.1: 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 3.2: 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.

Table 3.3: 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 3.4: 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.

3.6 Estimation

In this section we describe the estimation of equation (3.16), 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 3.6.1. 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 3.6.2. 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 the third issue is briefly described in subsection 3.6.3.

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:

$$y_{it} = \beta_0 + \beta_1 k_{it} + \beta_2 l_{it} + \beta_3 d(.,1)_{it} + \omega_{it} + \eta_{it} \qquad (3.20)$$

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. The authors' 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 further on, take a different approach by approximating the productivity shock ω_{it} using investment as a proxy variable. The authors 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 a lot of 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 (3.20) 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 like 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}.

In what follows 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}:

$$m_{it} = f(k_{it}, \omega_{it}, d(.,1)_{it}) \tag{3.21}$$

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

$$\omega_{it} = g(k_{it}, m_{it}, d(.,1)_{it}) \tag{3.22}$$

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}:

$$E(\eta_{it} \mid l_{it}, k_{it}, d(.,1)_{it}, m_{it}, \ldots, l_{i1}, k_{i1}, d(.,1)_{i1}, m_{i1}) = 0 \tag{3.23}$$

The dynamics of the unobserved productivity shock are restricted as:

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

$$= j(\omega_{it-1}) = j(g(k_{it-1}, m_{it-1}, d(.,1)_{it-1}) \tag{3.24}$$

For productivity innovations a_{it} we can write:

$$\omega_{it} = j(\omega_{it-1}) + a_{it} \tag{3.25}$$

where

$$E(a_{it} \mid k_{it}, d(.,1)_{it}, l_{it-1}, k_{it-1}, d(.,1)_{it-1}, m_{it-1}, \ldots) = 0 \tag{3.26}$$

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 (3.24) and (3.25) into the production function (3.20) we get:

$$y_{it} = \beta_0 + \beta_1 k_{it} + \beta_2 l_{it} + \beta_3 d(.,1)_{it} + j(g(k_{it-1}, m_{it-1}, d(.,1)_{it-1})) + u_{it} \tag{3.27}$$

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

$$E(u_{it} \mid k_{it}, d(.,1)_{it}, l_{it-1}, k_{it-1}, d(.,1)_{it-1}, m_{it-1}, \ldots) = 0 \tag{3.28}$$

Before estimating (3.27) 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 (3.25) can be rewritten as:

$$\omega_{it} = \psi + \omega_{it-1} + a_{it} \tag{3.29}$$

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

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

$$= h(k_{it}, m_{it}, d(.,1)_{it}, k_{it}m_{it}, k^2_{it}, m^2_{it}, k^2_{it}m_{it}, k_{it}m^2_{it}, k^3_{it}, m^3_{it}) \quad (3.30)$$

where h is a linear function. Using (3.29) and (3.30) we can rewrite (3.27) as:

$$y_{it} = (\beta_0 + \psi) + \beta_1 k_{it} + \beta_2 l_{it} + \beta_3 d(.,1)_{it} + g(k_{it-1}, m_{it-1}, d(.,1)_{it-1}) + u_{it} \quad (3.31)$$

Note that in (3.31) 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 (3.31) 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 (3.31), 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:

$$tfp_{it} = y_{it} - \beta_1 k_{it} - \beta_2 l_{it} \quad (3.32)$$

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.

3.6.2 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 (3.16), 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 (3.16). The multinomial probit approach is motivated by the unobserved ordering of trade strategies. As noted in section 3.3, 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:

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

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

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

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

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 (3.32), 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.

3.6.3 Estimation of the Equilibrium Sales Equation

Once tfp_{it} in (3.32) and the trade strategy probabilities have been fitted, all that remains is to estimate the equilibrium sales equation (3.16). 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 (3.17)–(3.19), as presented in Table 3.9 in the next section.

3.7 Results

Table 3.5 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

[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.

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 (3.30).

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 increases sharply after the correction. At the same time, the elasticity of labor stays 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 3.5 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 3.5. 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 stay 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 3.6. For the year-by-year estimates we refer the interested reader to Tables 3.A2–3.A5 in Appendix 2.

The coefficients on log real capital and log TFP in Table 3.6 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 3.7 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 3.8) discussed below and does not influence our main results regarding the exchange rate elasticities (Table 3.9).

By using the estimates of the sales equation in Table 3.7 we can derive estimates of some of the model's structural parameters. These are summarized and tested in Table 3.8. 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 (3.16).

Apart from the above structural parameters of the model we can use the estimates of the sales equation (3.16) 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 3.9 below. This is motivated by the fact that appreciation shocks usually get more attention in Czech economic news reports.

[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 (3.16) are correlated.

According to our results in Table 3.9, 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.

Table 3.5: 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 3.6: 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 3.7: 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 3.5, 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 3.5, column 8.

Table 3.8: 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 3.9: 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)	$-\rho 0(0,1) = -\rho(0,1) = -\rho 0(1,1)$	α_3	0.201***	0.227***
			(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.

3.8 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. At the same time, domestic sales are expected to rise by 0.2% and total sales to drop by 0.2% for the same sub-sample of firms. The above elasticities of export and total sales are roughly in line with our estimates on macro-level data.

We contributed to the literature on heterogeneous firms and trade by studying the impact of a hypothetical exchange rate shock to firm sales, a topic which has not been studied before in this context to our knowledge. 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.

Our research is also interesting from the point of view of estimating production functions. The findings concur with other studies regarding the importance of measurement error correction in capital. In particular, Galuščák and Lízal (2011) came to the same conclusion from a different Czech dataset. Moreover, our estimates imply that firm-specific fixed effects are likely to be endogenous in production functions. 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.

3.9 Appendix 1: Descriptive Statistics

Table 3.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.

3.10 Appendix 2: Year-by-Year Estimates of the Multinomial Probit

Tables 3.A2-3.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).

3.11 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 (3.16) 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 3.A9.

The exchange rate elasticities of exports and output on the macro-level are estimated by two simple approaches. The first, in (3.33) and (3.34), 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 (3.36) and (3.37) 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 (3.35), adapts equation (3.16) to the macro data. In particular, the export and import dummies in (3.16) 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 (3.18) and (3.19) on the firm level. Specifically, using the coefficients of (3.35), we can express w_x and w as in (3.38) and (3.39) below. Similarly to W_x and W, and following Table 3.7 in the Results section, w_x and w denote the elasticity of a 1% appreciation of the domestic currency.

$$B_0(L)X_t = b_{00} + b_1 REER_{t-1} + \delta_t \quad (3.33)$$

$$C_0(L)Y_t = c_{00} + c_1 REER_{t-1} + \varsigma_t \quad (3.34)$$

$$A_0(L)Y_t = a_{00} + a_1 XY_t + a_2 MY_t + a_3 TFP_{t-1} + \chi_t \quad (3.35)$$

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: $W_x = b_1$ (3.36)

Direct elasticity of output: $W = c_1$ (3.37)

Implied elasticity of exports: $w_x = a_3 - 1$ (3.38)

Implied elasticity of output: $w = a_3 + exp(-a_1) - 1$ (3.39)

Equations (3.33)–(3.35) 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 (3.33)–(3.35) 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 3.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 (3.33) and (3.34) 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 3.A8 lists the implied exchange rate elasticity estimates based on (3.35). 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

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

check. The implied elasticity estimates in the lower part of the first column are close to those obtained from the AR-X models (3.33) and (3.34). However, we cannot draw the same conclusion for the other combinations of data sources and estimation intervals presented in the remaining columns of Table 3.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 3.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 3.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 3.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 3.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.

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