The Impact of Ownership Structure on Firm and Market Performance

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Dedicated to Katharina

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Introduction

A key driver for the functioning of markets as an efficient allocation tool for goods and services is the ownership structure of the firms operating in the market. This thesis analyzes two distinct aspects of how the ownership structure influences firm behavior and market performance. First, ownership concentration impacts the strategic environment of competing firms, possibly resulting in a price distortion (Farrell and Shapiro, 1990). This aspect of ownership structure is the focus of chapters one and two, where ownership consolidation in the form of mergers are analyzed in an ex ante and ex post context, respectively. Second, the presence of institutional investors, such as banks, hedge funds, private equity firms, ect., within a firm can influence actions taken by the management and influences firm performance. The impact of institutional investors on innovation activity and firm productivity in terms of financial matters is analyzed in chapters three and four, respectively.

Ceteris paribus, common ownership of competing firms lessens price competition, resulting in an allocation distortion when goods are substitutes (Farrell and Shapiro, 1990). Compared to the separation of ownership, common ownership implies an additional trade-off when deciding on which price to set. In a nutshell, separated firms weigh off a positive demand expansion effect and an infra marginal loss, which occurs after a price decrease. The demand reduction of the competitors are not considered by each profit-maximizing firm. In a market with common ownership, the demand reduction of the other firm, following a price cut, is internalized, which gives each firm fewer incentives to lower their price. In a market with two firms, A and B, and common ownership after a price cut of firm A it gains consumers and loses infra marginally on the consumers that would also have bought at a higher price. In addition to this trade-off firm A also takes into account the loss in demand of firm B. This additional negative effect causes prices under common ownership to become higher than under separation. This anti-competitive effect is of particular interest to antitrust authorities when evaluating mergers and acquisitions. However, it is also well known that common ownership can exert a downward pressure on prices if synergy effects are present (Williamson, 1968; Farrell and Shapiro, 1990, 2001). The net effect of both these forces determines how efficiently a market performs under common ownership and separation. If synergy gains outweigh the market concentration effect, a merger is considered pro-competitive. One tool to assess the likely price effects of potential mergers is simulation, using a structural demand and supply model and performing a counterfactual analysis.

In chapter one, we extend the literature on merger simulation models by incorporating its potential synergy gains into structural econometric analysis. We present a three-step integrated approach. We estimate a structural demand and supply model, as in Bonnet and Dubois (2010). This model allows us to recover the marginal cost of each differentiated product. We then estimate potential efficiency gains using the Data Envelopment Analysis approach of Bogetoft and Wang (2005), and some assumptions about exogenous cost-shifters. In the last step, we simulate the new price equilibrium post-merger, taking into account synergy gains, and derive price and welfare effects. We use a homescan dataset of dairy dessert purchases in France, and show that for two of the three mergers considered, synergy gains could offset the upward pressure on prices. Some mergers could then be considered as not harmful for consumers.

Chapter two analyzes, ex post, the effects of a merger between a German supermarket chain and a soft discounter on consumer prices. We exploit geographic variation in prices within retail chains and brands and use a difference-in-differences estimator to compare regional markets with a change in market structure to a control group in unaffected markets. Our results indicate that both insiders and outsiders raised average prices after the merger, particularly in regions with high expected changes in retail concentration. In contrast, we estimate price declines in regions that did not experience a rise in concentration but were potentially affected by cost savings within the merged entity. We also provide evidence that remedies imposed by competition authorities were not sufficient to offset anti-competitive effects.

Institutional investors control a large portion of the equity of firms around the globe. Combined, they manage assets worth more than 90 trillion dollars OECD (2013). However, the impact of these investors is controversial. On the one hand they are accused of having a short-term view that is in contrast to the long term interests of the company. They have been publicly labeled as anonymous locusts that damage companies and disregard workers' interests.[1] On the other hand there is evidence that institutional investors act as a monitoring entity in firms, encouraging the management to take actions that have a positive impact on the company (Aghion et al., 2013; Bena et al., forthcoming). Empirical results suggest that finance is an important channel through which institutional investors can positively impact

[1]Interview of former chairman of the social democratic party, Franz Münterfehring, with Bild am Sonntag, April 17, 2005

firms. Asymmetric information in credit markets can lead to reductions in credit supply, leaving firms financially constrained (Stiglitz and Weiss, 1981). There is evidence that institutional investors can alleviate financial constraints and improve firm performance (Boucly et al., 2011; Amess et al., 2016; Agca and Mozumdar, 2008).

In chapter three, we analyze the relationship between institutional investors, innovation, and financing constraints. Building on the empirical framework of Aghion et al. (2013), we find that the effect of institutional ownership on innovation is concentrated in industries with high dependence on external finance and among firms which are a priori likely to be financially constrained. The complementarity between institutional ownership and competition, predicted by the original paper's theory where institutional investors increase innovation through a reduction in career risks, disappears once this heterogeneity is taken into account. We also provide evidence that the sensitivity of R&D investment to internal funds decreases with institutional ownership.

In chapter four, I analyze the relationship between foreign institutional investors and total factor productivity using a panel data set of European manufacturing firms. The results show a positive association of foreign institutional ownership on productivity. Using an instrumental variable estimator that exploits liquidity shocks of foreign institutional investors, caused by acquisitions of positions in the investors' portfolio, indicates the relationship is causal. The effect is more pronounced during the financial crisis and is driven by portfolio firms that are a priori more likely to be financially constrained. Furthermore, firms with a foreign institutional investor are less sensitive to cash flow shocks with respect to capital investments. The findings suggest that foreign institutional investors increase productivity by relaxing credit constraints.

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Chapter 1

An Empirical Analysis of Mergers: Efficiency Gains and Impact on Consumer Prices

Co-authored with Céline Bonnet

1.1 Introduction

The merger of firms play an important role in an economy and are addressed by public policy. Annually, the Federal Trade Commission (FTC) and the Department of Justice (DOJ) reviews mergers valued at over a trillion U.S. dollars in total. During 2000 and 2010, the DOJ triggered around 1,700 investigations for merger cases in order to assess possible anti-competitive effects (Farrell and Shapiro, 2010). However, most mergers are ruled as being unproblematic, which suggests that most mergers are neutral or pro-competitive and only a small fraction are revised. The competition authorities first use market concentration tests, such as the Herfindahl-Hirschman-Index (HHI), to gauge whether or not prospective mergers would produce a firm of such a magnitude that it would adversely impact social welfare. The advantage of this test is that data on market shares can be observed, and the HHI and its change post merger can be easily computed. The merger guidelines of the European Commission[1] state that mergers in markets with a post-merger HHI below 1,000 are generally unproblematic. Mergers with a post-merger HHI between 1,000 and 2,000, and a change in concentration below 250, are also considered to be unproblematic. When the post-merger HHI is higher than 2,000, the threshold for the critical change is 150. However, the HHI index is a very crude measure for likely competitive effects post merger and does not capture other important factors that that may increase or decrease consumer welfare. Among other factors, the particular focus of this paper is the strategic reaction of all firms in the industry and the synergy gains induced by a merger.

In the case of horizontal mergers, there are two opposing forces in terms of price. The higher market concentration exerts an upward pressure on prices as the merged firm has more market power. This in turn allows competitors outside the merger to increase their prices, such that in equilibrium the whole industry raises their prices. On the other hand, synergy gains in the form of lower marginal costs lead to a downward pressure on prices after a merger. The question is what is the net effect of these two forces? In Farrell and Shapiro (1990), one of the main results is that mergers without synergies always lead to price increases and consumers are worse off.

Since the mid 90's, merger simulation models based on the New Empirical In-

[1]The french competition authority uses the same tests to determine if a merger is likely to be anti competitive

dustrial Organization literature have been increasingly used in antitrust cases in the US and Europe as a complementary tool, in addition to classical methods that are based on market concentration. These simulations consist of describing consumer behavior in the market in terms of substitution patterns and the strategic behavior of firms, and then taking into account the strategic behavior of all competitors in the industry. Nevo (2000) estimates a random coefficient logit model in order to analyze merger effects in the ready-to-eat cereal industry. Ivaldi and Verboven (2005) use a nested logit model to represent the customers' preferences on the truck market and assess the effects of the merger between Volvo and Scania. In their study, they compare the merger effects without synergies and with a hypothetical 5% decrease in marginal costs. Bonnet and Dubois (2010) conduct a counter-factual analysis of a de-merger between Nestlé and Perrier in a structural model of vertical relationships between bottled water manufacturers and retail chains. Fan (2013) assesses the merger effects in the newspaper market allowing for adjustments in product characteristics as well as prices. Mazzeo et al. (2014) also look at merger effects when firms can reposition their products post merger. Earlier work on merger analysis is given, for example, by Baker and Bresnahan (1985), Berry and Pakes (1993), Hausman and Leonard (2005), Werden and Froeb (1994) and Werden and Froeb (1996). Even though the development of simulation models is still young, their importance is likely to grow in the future as computational power and the availability of data steadily increases. However, these models assume that either the cost structure does not change post merger and then there is no synergy effect, or ad hoc synergy gains as a 5% decrease in marginal cost. Without synergies, we only capture the force that exerts the upward pressure on prices due to a higher concentration, which implies that the results can be interpreted as a maximal benchmark in this case (Budzinski and Ruhmer (2009)). Ashenfelter et al. (2015) show empirical evidence that synergies can offset price increases, which supports our findings. They find that a merger in the beer industry actually produced enough synergy gains to offset the effect due to increased market concentration. They show that prices increase by 2% without synergies and that synergy countered this effect on average.

There is only few literature on estimating cost efficiencies for counter factual horizontal mergers. Jeziorski (2014) analyzes the merger wave in the US radio market and presents a method to estimate fix cost efficiencies for mergers. Another approach is presented in Mazzeo et al. (2014) that also analyzes potential fix cost efficiencies of mergers. Both papers involve product repositioning post merger and savings are realized through fix costs. We propose a new approach to extend the merger simulation literature taking into account marginal costs savings as a source of synergy gains. The method is computationally inexpensive and can be applied to any market where mergers are likely to reduce marginal costs. This approach consists of a three-step method. First, we develop an empirical model of demand and supply as in Villas-Boas (2007) or Bonnet and Dubois (2010).[2] For the supply side we consider a vertical structure as in Rey and Verge (2010), extended to multiple upstream and downstream firms. The retailers on the downstream level engage in price competition for consumers and are supplied by the upstream firms. It is assumed that two-part tariff contracts are used in the vertical relationship in order to avoid the double marginalization problem and maximize profits in the vertical chain. The demand side is estimated using a random coefficient logit model, which ensures the flexible substitution patterns of consumers. Using an integrated structural econometric model which takes into account the consumer substitution patterns and the strategic reaction of firms in the market considered, we are able to recover marginal costs for each product sold in the market. Using exogenous cost shifters, we are able to estimate the impact of some inputs on the marginal costs and then assess the quantity of inputs needed to produce one additional unit of product. The second step uses the Data Envelopment Analysis (DEA) method to estimate the potential efficiency gains of mergers following Bogetoft and Wang (2005). We then obtain an estimated amount of marginal cost savings for the merger. Finally, integrating the change in the structure of the industry induced by the merger and the change in marginal cost for each product, we are able, in a third step, to assess the total effect of the merger on consumer prices.

We apply this integrated approach to the French dairy dessert market. This sector is composed of six main manufacturers (Danone, Nestlé, Senoble-Triballat, Yoplait, Andros, Sojasun and Mom) which sell national brand products and some small to medium firms that sell private label products. We estimate the synergy gains for all 78 possible bilateral mergers. The derived synergy gains from the Data Envelopment Analysis focus on merger gains that arise due to economies of scope, and leaves out other potential sources of gain, such as scale economies. There is a large heterogeneity of potential merger gains which shows that the 5% ad hoc rule

[2] Also see Rosse (1970), Bresnahan (1989), Berry (1994), Goldberg (1995), Berry et al. (1995), Slade (2004), Verboven (1996), Nevo (1998), Nevo (2000), Nevo (2001) and Ivaldi and Verboven (2005)

can be very misleading. Considering all possible mergers, we find that around 44% of the mergers are expected to produce no synergies at all, while 18% of the mergers are expected to produce synergies higher than 5%. The remaining 38% produce synergies lower than 5%. Considering all bilateral mergers, the average marginal cost savings are 2.55% which would then be the best approximation ex ante for merger gains in this industry.

We choose to simulate the new price equilibrium for three selected mergers with low and high cost savings and for low and high changes in market concentration. To disentangle the effect of synergy gains and market concentration, we compare four scenarios for the three selected mergers. First, we simulate the mergers without any reduction in marginal costs, taking into account only the industry change; second we estimate the merger without the structural change in the industry considering only the lower marginal costs; and third, we take into account both the new industry structure and the estimated reductions in marginal costs of the merged entity. Finally, in order to compare with the common analyses of merger simulations, we simulate the 5% synergy gain assumption.

As expected, without synergy gains, all three mergers lead to an increase in industry prices, a decrease in consumer surplus and an increase in industry profits. In the case when we consider only the reduced marginal costs of the merging firms but not the change of the industry structure, prices decrease for both merging firms and outside firms. Firms that experience the savings always increase profits, while outside firms always lose in this scenario. Not surprisingly, consumer surplus increases.

Taking into account the structural changes of the industry and the reduction of the marginal cost delivers the net effect on the market of the two opposing forces. The three mergers we consider have a difference in HHI and a cost reduction of 390 and 7.84%, 195 and 1.88%, and 67 and 3.37%, respectively. The first merger represents a case of two main players in the market that have large potential savings. The second merger is a case where a major player merges with a private label producer. The last merger is a case of two private label manufacturers. We find that the downward pricing pressure can outweigh the upward pricing pressure, as is the case in the first and third merger. Our results suggest that merged firms may employ an asymmetric pricing strategy post merger in order to shift consumers from one manufacturer to the other. We find that the 5% ad hoc cost saving rule is not appropriate as a rule of thumb for the considered mergers as it does not reliably capture market outcomes post merger. Alongside the methodological contribution, we also aim to disentangle the effects of synergy gains and concentration effects. For most market outcomes, such as changes in prices, industry profits and consumer surplus, our results suggest that concentration and synergies have an equal weight on the direction of the effect. By comparing the net effect to the two benchmark cases we see that these effects lie in the middle of both extremes. However, this is not true for an important outcome; that is, the changes in the profits of the merged entities. We find that the changes in profits are practically only driven by synergy gains. This is important as the changes in profits are the main incentive for managers to merge. This suggests it is not the expected profit increase due to concentration but rather to synergy gains that creates merger incentives.

In summary, we make several contributions. First, we propose a framework to estimate potential synergy gains of horizontal mergers. Second, we show that in the French dairy dessert industry, the 5% ad hoc rule overstates the pro-competitive effects of mergers. Third, we show that firms may use asymmetric pricing strategies post merger. Fourth, while for changes in consumer surplus post merger, upward and downward pricing pressure has a similar impact on the final effect, firms' incentives to merge are only driven by potential cost savings.

The rest of the article is organized as follows. Section 2 presents the facts of the considered market and summary statistics. Section 3 describes the methodological approach containing the supply and demand model, the DEA method and how it is used to estimate potential merger gains, and the simulation method that allows us to estimate the new price equilibrium. Section 4 shows the empirical findings, and last, Section 5 concludes.

1.2 The French dairy dessert market

As an application, we focus on the French dairy dessert market for two reasons. First, this market is composed of oligopolies with market power, in which manufacturers adjust margins when they face cost shocks. Second, dairy desserts are mainly composed of milk which simplifies the specification of the cost function in this sector.

We use home scan data on food products provided by the society Kantar TNS WorldPanel, 2011. These data include a variety of purchase characteristics, such as quantity, price, retailer, brand, and product characteristics. We consider the market of desserts to be composed of 30% of purchases of dairy desserts and 70% of purchases of other products, that we will call the "outside good".[3] The data set contains more than 2.5 million purchases over the year 2011. We consider seven retailers: five main retailers and two aggregates (one for the hard discounters and one for the remaining hypermarkets and supermarkets). We also consider six manufacturers of national brands and seven manufacturers of store brands, one for each retailer. Table 1 reports market shares by producer and category. Manufacturers 7-13, which represent the private label products, have a combined market share of around 41%. The market shares for national brand manufacturers vary from 1% to 24%. Table 2 presents average prices by manufacturer and category. On average, the private label products have the lowest retail prices followed by the three largest manufacturers 1 to 3. The small firms 4, 5 and 6 charge the highest retail prices on average. By category, yoghurts have the lowest retail prices followed by desserts and petit suisse, respectively.

We consider 26 brands including an aggregate brand for private labels for each of the three categories. As products are defined by the combination of brands and retailers, we get 162 products for each month.

1.3 Methodology

Here we introduce the methodology used for estimating the equilibrium model, efficiency gains of potential mergers, and counterfactual analysis. First, we introduce the demand and supply framework of Bonnet and Dubois (2010) that allows us to recover the marginal cost of each product in the market. Second, we show how Bogetoft and Wang (2005) use DEA to estimate efficiency gains of mergers and how we incorporate this method into the framework of Bonnet and Dubois (2010). Using the estimated marginal costs and exogenous prices of inputs, we are then able to recover the quantity of inputs needed to produce each product at each time period. Finally, we present the merger simulation method that also accounts for the new industry structure and synergy gains.

[3] This outside option is composed of all other desserts, such as fruits, pastries, and ice creams.

1.3.1 The equilibrium model

In this section, we introduce the demand and supply specification and present the identification and estimation strategy. For the demand, we use a random coefficient logit model as it allows for flexible substitution patterns for consumers between alternatives. Compared to the standard logit model, this leads to more realistic estimates of own- and cross-price elasticities, which in turn gives better margins and thus better estimates for marginal costs. For the supply model, we use two-part tariff contracts between manufacturers and retailers, as in Bonnet and Dubois (2010).

Demand model and identification

We use a random utility approach and in particular a random coefficient logit model, as in McFadden and Train (2000). We define the indirect utility function of consumer i of buying product j at time t as:

$$V_{ijt} = \beta_{b(j)} + \beta_{r(j)} + \alpha_i p_{jt} + \xi_{jt} + \epsilon_{ijt}, \tag{1.1}$$

where $\beta_{b(j)}$ and $\beta_{r(j)}$ are time invariant brand and retailer fixed effects, respectively, p_{jt} is the price of product j at time t, ϵ_{ijt} is an unobserved individual error term that is distributed according an extreme value distribution of type I, and ξ_{jt} is also unobserved by the econometrician and represents changes in product characteristics over time. We allow for consumer heterogeneity for the disutility of the price specifying the price coefficient α_i as follows:

$$\alpha_i = \alpha + \sigma v_i, \tag{1.2}$$

where α is the average price sensitivity, v_i follows a normal distribution that represents the deviation to the average price sensitivity and σ is the degree of heterogeneity.

We allow for an outside option, that is, the consumer can choose another alternative from amongst the J products of the choice set. The utility of the outside good is normalized to zero, which means $U_{i0t} = \epsilon_{i0t}$.

The mean utility is defined as $\delta_{jt} = \beta_{b(j)} + \beta_{r(j)} - \alpha p_{jt} + \xi_{jt}$ with a deviation defined as $\nu_{ijt} = -\sigma v_i p_{jt}$, such that we get:

$$V_{ijt} = \delta_{jt} + \nu_{ijt} + \epsilon_{ijt}. \tag{1.3}$$

The individual probability of buying the product j takes the logit formula as follows:

$$s_{ijt} \equiv \frac{exp\left(\delta_{jt} + \nu_{ijt}\right)}{1 + \sum_{k=1}^{J_t} exp\left(\delta_{kt} + \nu_{ikt}\right)}. \tag{1.4}$$

The aggregated market share of product j is thus given by:

$$s_{jt} = \int_{A_{jt}} s_{ijt}\, \phi\left(v_i\right) dv_i, \tag{1.5}$$

where A_{jt} is the set of consumers buying the product j at time t and ϕ is the density of the normal distribution.
The own- and cross-price elasticities are given by:

$$\frac{\partial s_{jt}}{\partial p_{kt}} \frac{p_{kt}}{s_{jt}} = \begin{cases} -\frac{p_{kt}}{s_{jt}} \displaystyle\int \alpha_i s_{ijt} \left(1 - s_{ijt}\right)\ \phi\left(v_i\right) dv_i & if \quad j = k \\ \frac{p_{kt}}{s_{jt}} \displaystyle\int \alpha_i s_{ijt} s_{ikt}\, \phi\left(v_i\right) dv_i & otherwise. \end{cases} \tag{1.6}$$

The unobserved term ξ_{jt} is likely to be correlated with price as it captures, for example the advertising expenses of the manufacturers. Advertising influences purchasing decisions by consumers, and as it is costly, it will certainly also affect the price. Failing to control for this unobserved product characteristic results in a correlation between price and the error term of the demand equation, leading to a biased estimate of the price coefficient. In order to deal with this issue, we use the two-stage residual inclusion approach (Terza et al. (2008); Petrin and Train (2010)). In the first step, we regress prices on a set of instrumental variables and exogenous variables of the demand equation. As instruments, we use input prices, noted W_{jt}:

$$p_{jt} = W_{jt}\gamma + \theta_{b(j)} + \theta_{r(j)} + \eta_{jt}, \tag{1.7}$$

where $\theta_{b(j)}$ and $\theta_{r(j)}$ are brand and retailer fixed effects, respectively. The error term η_{jt} captures the effect of the omitted variables, such as advertising costs. If we include this error term in the mean utility of the demand equation as regressor, it captures the effect of the unobserved characteristics on price. This means the error term of the demand equation, $\zeta_{jt} = \xi_{jt} - \lambda\hat{\eta}_{jt}$, will not be correlated with the price. We can then write the mean utility as:

$$\delta_{jt} = \beta_{b(j)} + \beta_{r(j)} - \alpha p_{jt} + \lambda\hat{\eta}_{jt} + \zeta_{jt}. \tag{1.8}$$

Supply model

This section presents the theoretical model we use for the supply side. Manufacturers and retailers engage in a vertical relationship. We know that linear pricing, where the downstream firm pays a per unit price to the manufacturer, leads to a double marginalization problem, and profits in this chain are not maximized. This is why the parties often use two-part tariffs. We then use the same empirical framework of Bonnet and Dubois (2010) to model the vertical relationships in the dairy dessert market. We also assume that manufacturers impose the final prices on the retailers. Bonnet et al. (2015) show that these kind of contracts are preferred to linear contracts or two-part tariff contracts without resale price maintenance in the dairy dessert market. Other empirical studies suggest that this contract is used in other markets as well (Bonnet and Réquillart (2013); Bonnet and Dubois (2010); Bonnet et al. (2015)). Furthermore, it is assumed that manufacturers have all the market power with respect to the retailers. The game between manufacturers and retailers is then described in the following:

Stage 1: Manufacturers simultaneously propose two-part tariff contracts to the retailers. It is assumed that those contracts are public[4] and consist of a per unit wholesale price, a fixed fee, and the final price of the products.

Stage 2: After observing the offers, the retailers simultaneously accept or reject them. In case of the rejection of a retailer, they earn an exogenous outside option. If all retailers accept, demand and contracts are satisfied.

Rey and Verge (2010) prove the existence of a continuum of equilibrium in which consumer prices are decreasing in wholesale prices.

Let there be J different products defined by the cartesian product of brand and retailer sets. Let there be R retailers, and S_r is the set of products sold by

[4] Following the argument of Bonnet and Dubois (2010), this assumption is justified by the non-discrimination laws of comparable services. They argue that in the case of resale price maintenance only the offered retail prices are essential for the decision-making of the retailers.

retailer r. There are F manufacturers and let G_f be the set of products produced by manufacturer f. F_j is the fixed fee paid by the retailer for selling product j, w_j is the according wholesale price and p_j is the final retail price. Let $s_j(p)$ be the market share of product j, M is the total market size, and μ_j and c_j are the constant marginal costs of production and distribution of the product j, respectively.
The profit of the retailer r is given by:

$$\Pi^r = \sum_{j \in S_r} [M(p_j - w_j - c_j) s_j(p) - F_j], \tag{1.9}$$

and the profit of the manufacturer f is:

$$\Pi^f = \sum_{k \in G_f} [M(w_k - \mu_k) s_k(p) + F_k], \tag{1.10}$$

subject to the retailers' participation constraints: $\forall\ r = 1, ..., R,\ \Pi^r \geq \overline{\Pi}^r$. As already explained, manufacturers make take-it or leave-it offers to retailers. However, they have to respect the participation constraint of the retailers. Let the outside option $\overline{\Pi}^r$ of the retailer be a constant that can be normalized to zero, such that we have for the participation constraint $\Pi^r \geq 0$. As Rey and Verge (2010) show, this participation constraint is binding and can be substituted in (1.10) which gives the following expression:

$$\Pi^f = \sum_{k \in G_f} [M(p_k - \mu_k - c_k) s_k(p)] + \sum_{k \notin G_f} [M(p_k - w_k - c_k) s_k(p)] - \sum_{j \notin G_f} F_j. \tag{1.11}$$

Rey and Verge (2010) point out that a manufacturer internalizes the margin of the whole vertical chain of its own products, but also internalizes the retail margin of its competitors. We see that this profit expression does not depend on the franchise fees of the manufacturer f. As the manufacturers simultaneously propose the contracts to the retailers, the wholesale prices and the fixed franchise fees of the other manufacturers do not affect the terms of the contracts offered by the manufacturer f. Moreover, when resale price maintenance is introduced, the wholesale prices do not affect the profit of the manufacturer as it can conquer the retail profits via the franchise fee. However, it does influence the behavior of its competitors. The optimization problem of f then becomes:

$$\max_{p_k \in G_f} \sum_{k \in G_f} \left[M \left(p_k - \mu_k - c_k \right) s_k \left(p \right) \right] + \sum_{k \notin G_f} \left[M \left(p_k^* - w_k^* - c_k \right) s_k \left(p \right) \right]. \tag{1.12}$$

In the absence of any additional restrictions, we are not able to separately identify the wholesale margins $w_k - \mu_k$ and the retail margins $p_k - w_k - c_k$. As suggested by Bonnet et al. (2015), we will assume that retail margins are equal to zero in the French dairy dessert market. The first order conditions for the manufacturer f under this condition are given by:

$$s_j \left(p \right) + \sum_{k \in G_f} \left(p_k - \mu_k - c_k \right) \frac{\partial s_k \left(p \right)}{\partial p_j} = 0 \quad \forall j \in G_f. \tag{1.13}$$

The first order conditions allow us to recover marginal costs mc_{jt} as the sum of costs of production and distribution. Note that the identified margin $p_k - \mu_k - c_k$ is equal to $w_k - \mu_k$ under the assumption of zero retail margins. It is useful to denote the identified margins in matrix notation as follows:

$$\hat{\Gamma}_f + \hat{\gamma}_f = - \left(I_f S_p I_f \right)^{-1} I_f s \left(p \right) \quad \forall f \in F. \tag{1.14}$$

The left-hand side is the sum of wholesale ($\hat{\Gamma}_f$) and retail ($\hat{\gamma}_f$) margins.[5] The margins are identified as a function of estimated demand parameters that are included in S_p, that is, the matrix of first derivatives of market shares with respect to price. I_f is the diagonal ownership matrix of firm f which takes the value 1 if the product belongs to the firm f and 0 otherwise. $s\left(p\right)$ is the vector of estimated market shares.

Using observed prices and estimated margins, we are able to recover marginal costs which are then used to estimate inputs in the next step.

1.3.2 Merger simulation

When assessing the implications of a merger, two main factors are in opposition. First, the change in industry implies an upward pressure on price, as the market power of the merged firm increases. Competitors also benefit from the decreased competitive pressure by increasing their prices. The second factor is the change of

[5]Note that in our selected equilibrium, the latter one is equal to zero.

the cost structure of the merged firms. Synergies may reduce marginal production costs that firms can partially or completely transmit to final retail prices. Competitors have to follow and cut prices to prevent their customers from switching. The question is, what is the net effect on the price of these two opposing forces?

In this section, we present the methodology used to estimate cost savings from a potential merger, on the one hand, and the methodology used to evaluate the price effect of a merger — taking into account the estimated cost saving and the change of the industry — on the other hand. Section 3.2.1 describes the source of potential cost savings. In Sections 3.2.2 and 3.2.3, we present the methodology on how we estimate the potential savings of production inputs. The method presented in Section 3.2.3 requires data on the inputs and the outputs of manufacturers. In Section 3.2.2, we show how we obtain these quantities from our data. Section 3.2.4 shows how we translate the input savings into marginal cost reductions, and finally, Section 3.2.5 describes how we simulate the new price equilibrium post merger and how we assess the effect on consumer surplus.

Synergies

Cost efficiencies are a major argument used by firms to justify a proposed merger in front of the regulatory authorities. Gugler et al. (2003) showed that synergies are present in about one third of the mergers. In this section, we discuss where synergies stem from in general, the relevant merger gains to focus on and we examine their suitability in the French dairy dessert industry.

Farrell and Shapiro (1990) describe three sources of merger gains. First, the merged firms can distribute their output across the participating firms, moving production to the more efficient production sites. This would not change the firms' knowledge or capital intensity. The changes in marginal costs and the associated gains are due to a reallocation to the production site with lower marginal costs pre-merger. Second, firms can shift capital from one firm to another in order to operate at a higher scale. Farrell and Shapiro (2001) argue that merger benefits that are due to economies of scale cannot be used to justify a merger in most cases. This is because if firms can reach a higher scale of production, and thus benefit from lower production costs without a merger, the benefits of a merger can be realized without an increase in market concentration which would be potentially harmful for consumers. They further explain that in the general case, market pressure should force firms to produce at a higher scale if possible, making a merger unnecessary.

Third, firms can learn from each other. This involves the sharing of knowledge and management skills, or taking advantage of complementary patents. This could lead to a change in marginal cost after a merger. In this case, the marginal cost reduction is derived from economies of scope or harmony effects (Bogetoft and Wang (2005)).

An important question in this paper is identifying the source of synergies which is most likely to occur in the French dairy dessert industry. The main input ingredient of any yoghurt or other dairy desserts is raw milk. In the production process, there are multiple highly-automated steps that the raw milk goes through in order to transform it into an input that can be used in the final product. The first steps (clarification and standardization) in the production process usually involve multiple rounds in centrifuges and cooling and heating phases in order to ensure a homogeneous input that meets certain requirements, such as fat content. Depending on the final product, different steps then follow which are automated and firm specific. Each company is likely to be efficient in some of the steps in the production process that are crucial for their specific products. Certain products require certain procedures in order to obtain the desired consistency and water content, which makes each firm an expert in the method it uses in the production process. After a merger, this firm-specific expertise is shared among the merging firms, such that we expect that in this industry, technological transfer and management skills are the main source of merger gains (Brush et al. (2011)). Another important factor is the number of people involved in the production process. A firms' technical sophistication should also be reflected in the number of workers required for production. A more technically-advanced firm requires fewer workers to produce a certain amount of output; that is, better management can make use of a workforce more efficiently. After a merger, technical transfer and introduced management skills should also carry over to the less efficient firm in terms of manpower efficiency.

Inputs and outputs

After obtaining marginal costs mc_{jt} from the equilibrium model, in order to calculate synergy gains, we need definitions for inputs $k \in K$ and outputs. As we have to make outputs of arbitrary firms comparable, it is convenient to define a product by its category $c \in C$. We then distinguish the outputs between the three different categories: yoghurts, fromage frais and petit suisse, plus other dairy desserts. This means that each firm $f \in F$ can produce three different kinds of products. The total output of firm f of product category c is given by its total quantity of products

$y_{fc} = Ms_{fc}$ in this category, where s_{fc} is the market share of firm f in the product category c and M is the total market size. We assume that production technologies in each category across firms differ in efficiency. We estimate inputs by regressing marginal costs on input prices g_{kt} and we assume that the marginal cost of product j in period t has the following specification:

$$mc_{jt} = \sum_{f \in F} \beta_{f(j)c(j)} + \sum_{r \in R} \beta_{r(j)} + \sum_{f \in F} \sum_{c \in C} \sum_{k \in K} \gamma_{f(j)c(j)k} g_{kt} + \epsilon_{jt}. \tag{1.15}$$

As marginal costs are the costs of producing one unit,[6] the estimated coefficients $\gamma_{f(j)c(j)k}$ can be interpreted as the amount of input k needed to produce one unit. $\beta_{f(j)c(j)}$ are fixed effects for each firm and category that capture unobserved cost shifters that vary across firms and product categories. $\beta_{r(j)}$ are retailer fixed effects and capture retailer specific costs, such as costs of warehousing and distribution. The total amount of input k for product category c of firm f is given by $x_{fck} = Ms_{fc}\hat{\gamma}_{fck}$. A firm f is characterized by its input and output vector x^f and y^f, respectively.

Estimating potential merger gains with DEA

We now introduce the method that allows us to estimate the potential gains ex ante using Data Envelopment Analysis (DEA).[7] Following Bogetoft and Wang (2005), the production possibility set $T(x, y)$ is defined as:

$$T = \left\{ (x, y) \in \mathbb{R}_+^{K+C} | \text{x can produce y} \right\}. \tag{1.16}$$

In the following, we introduce how synergy gains are estimated in this framework. First, we introduce the "potential overall gains" of merging, which is a maximal

[6]We assume a constant return to scale production technology. Under this assumption, we can use the marginal costs as average costs. (cf Footnote 11 in Subsection 3.2.3 for a justification of this assumption)

[7]DEA is a non-parametric method for production frontier estimation and is used to measure the efficiency of a single firm within an industry. The basic idea of the efficiency measure is linked to Farrell (1957) who argues that the efficiency of a firm is assessed by its distance to the production possibility set of the industry. The method wraps a tight fitting hull around the input/output data of the entire industry, such that it satisfies certain assumptions that vary, depending on the specification of the technology. The efficiency of a firm is then measured by its distance to this hull, also called frontier. This method can be applied to any d production plants, firms, and even aggregated industries (Balk, 2003). For a more fundamental introduction, see for example, "Introduction to Data Envelopment Analysis and Its Uses" by Cooper et al. (2005).

benchmark of potential savings. Then, in order to receive more realistic potential savings, "adjusted overall gains" are introduced, which control for the individual inefficiency of the merging firms. In the third step, the concept of "input slacks" is introduced, allowing us to calculate additional savings that are not captured in the previous step. At the end of this section, we provide three examples to illustrate the importance of the assessment of adjusted overall gains and input slacks.

Potential overall gains

In order to measure the potential synergy gains of a merger, the inputs and outputs of the merged entity are pooled together, pretending that it is in fact one firm, and then the distance of this artificial firm to the frontier is measured. There are several ways to conduct this projection. We use the input-oriented view, as this can be converted to cost savings. With this approach, the efficiency of firm f is measured by how much we can reduce the inputs of this firm and produce the same amount of outputs in comparison to the industry frontier. Let us define the set of firms that merge as U. Pooling the inputs and outputs of all $f \in U$ and projecting it on the frontier gives the "potential overall gains" from merging, and is described by the following program:

$$E^U = Min \left\{ E \in \mathbb{R} \,\middle|\, \left(E * \sum_{f \in U} x^f, \sum_{f \in U} y^f \right) \in T \right\}. \tag{1.17}$$

This translates into the following optimization problem:

$$\begin{aligned} \min_{E,\lambda} E \quad & s.t. \\ E * \left[\sum_{f \in U} x^f \right] \quad & \geq \quad \sum_{f \in F} \lambda^f x^f \\ \sum_{f \in U} y^f \quad & \leq \quad \sum_{f \in F} \lambda^f y^f \\ \lambda \quad & \geq \quad 0 \end{aligned} \tag{1.18}$$

As E is a scalar, it reflects the proportional reduction of inputs that is possible compared to the efficient industry frontier. The condition $\lambda \geq 0$ comes from the constant return to scale assumption.[8] The measure E^U simply represents the inefficiency of the combined inputs and outputs of the merging firms U. It allows us to compute the amount of reduction (1-E^U) in percentage of the inputs used by the merged firms to reach the industry frontier. Note that E^U captures more than the synergy effects that are due to economies of scope as it also contains the individual inefficiency of the individual merging firms (Bogetoft and Wang (2005)).

Adjusted Overall Gains

As we aim to extract the effects that are due to learning from amongst the merging firms, we have to remove the effect due to the technical inefficiency of the individual firms. Following Bogetoft and Wang (2005), we decompose the "potential overall gains from merging" E^U into the "adjusted overall gains" E^{U^*} and individual inefficiency T^U in such a way that we have:

$$E^U = E^{U^*} * T^U. \tag{1.19}$$

Bogetoft and Wang (2005) propose to first project the inputs of the individual firms on the frontier and then use these adjusted inputs in the merger analysis. This way individual inefficiency is taken care of, and the resulting efficiency measure of the pooled adjusted inputs only reflect the synergy effects due to learning or economies of scope. E^{U^*} is given by:

$$E^{U^*} = Min\left\{E \in \mathbb{R} \,\middle|\, \left(E * \sum_{f\in U} E^f x^f, \sum_{f\in U} y^f\right) \in T\right\}. \tag{1.20}$$

We see that the only difference between E^{U^*} and E^U is that we multiply the input vector of the individual firm f, that is part of the merger U, by its efficiency

[8] DEA can be performed under different assumptions in terms of the underlying production technology T of the considered sector. The returns to scale property plays a crucial role as the shape of the estimated frontier depends on this property. Potential merger gains are calculated using the distance of the combined firm's inputs and outputs to the frontier. This means the merger gains are dependent on the shape of the frontier. Thus, it is crucial to provide a good justification for the chosen technology. As explained in Section 3.2.1, the source of merger gains we focus on is learning and technology sharing. Following the argument of Farrell and Shapiro (2001), we rule out increasing returns to scale. Furthermore, following Bogetoft and Wang (2005), we assume that after a merger, firms could further operate as two single firms and simply coordinate prices without further integration. This rules out decreasing returns to scale. In order to capture synergies derived from economies of scope, we will then use a constant returns to scale (CRS) production technology.

score E^f obtained from a first step where we projected the single firms on the frontier.[9]

Input slacks

Thus far, we have been considering a linear reduction in inputs, which is a proportional decrease of all inputs by $\left(1 - E^{U^*}\right) * 100\%$. However, it is likely that after this proportional reduction, there is still the possibility of input specific reductions. Ji and Lee (2010) refer to this as "input slacks". Bogetoft and Wang (2005) mention this possibility but do not implement it. Here, we make use of this option and add the input slacks to allow for nonlinear reductions in inputs. Intuitively, the achieved savings in this case are at least as high as in the linear case, because the input slacks are on top of the linear part.

We define the input slack for input k of the merged entity U as S_k^U. Let x_k^U be the pooled total amount of input k for the entity U. Then the reduced input k post merger is given by:

$$E^{U^*} x_k^U - S_k^U. \tag{1.21}$$

We want to express the savings of input k of U in a single score E_k^U similar to the efficiency scores obtained from the DEA but input specific. Following Cooper et al. (2005), this is given by:

$$E_k^U * x_k^U = E^{U^*} * x_k^U - S_k^U. \tag{1.22}$$

The first term of the right-hand side $E^{U^*} x_k^U$ is the linear part of the reduction. In addition to this, we subtract the input specific reduction S_k^U. We want to express the total reduction of each input as an input specific score E_k^U. We have all information such that we can solve for E_k^U, as follows:

$$E_k^U = \frac{E^{U^*} * x_k^U - S_k^U}{x_k^U}. \tag{1.23}$$

[9] This first step is a standard DEA with a CRS technology using all initial firms. From this, the efficiency scores E^f are obtained and are then used to adjust the inputs of the merging firms.

Examples

To illustrate the three different parts that have been formally introduced, consider the following examples. There are two firms, A and B, neither of which are located on the frontier, and that produce both one product y with two inputs x_1 and x_2. Let us assume that at the efficient industry frontier, in order to produce $y = 10$ outputs, $x_1 = 4$ and $x_2 = 4$ inputs are required. Let us distinguish three different settings of this market to illustrate the three parts shown before: a) firms use the same technology; b) firms use different technology symmetrically; and c) firms use different technology asymmetrically.[10] In a), both firms are equally efficient. This means that they need the same number of inputs to produce the same number of outputs, that is, $y = 5$, $x_1^A = 5$ and $x_2^A = 5$ and $x_1^B = 5$ and $x_2^B = 5$. In this setting, both firms use the same production technology, and we would not expect any merger gains that are due to economies of scope or learning, as the firms are equal. However, in this setting, the measure E^U would predict merger gains as follows. When inputs and outputs are pooled, we have $y = 10$, $x_1^U = 10$ and $x_2^U = 10$. We can reduce both inputs by 60% compared to the frontier because we then hit the frontier in both input dimensions. This now means that we have $E^U = 0.4$ because $0.4 * x_1^U = 4$. This example shows that two equal firms would produce synergy gains using E^U. These synergies are due to both firms' individual inefficiency pre-merger, as the example illustrates.

Consider now the example in setting b) where neither firm is located on the frontier but use different production technologies. Let us now assume that firm A uses $x_1^A = 4$ and $x_2^A = 6$ and firm B uses $x_1^B = 6$ and $x_2^B = 4$ to produce $y = 5$ outputs each. Let the efficient frontier be $y = 5$ and $x_1 = 2$ and $x_2 = 2$. We see that both firms are inefficient compared to the frontier, but firm A has an advantage in input 1 compared to firm B and firm B has an advantage in input 2 compared to firm A. The program E^U (without removing individual inefficiency) would produce the following savings. Combining both firms' inputs and outputs we have $y = 10$, $x_1^U = 10$ and $x_2^U = 10$. We can reduce both inputs by 60% because we then hit the frontier in both dimensions. This implies that $E^U = 0.4$. This again contains individual inefficiency plus the gains that are due to learning. If we now project the individual inputs (without pooling) on the frontier we get $E^A = 0.5$, $E^A * x_1^A = 2$

[10]Table 7 in the Appendix summarizes the results of the three examples.

and $E^A * x_2^A = 3$ for firm A and $E^B = 0.5$, $E^B * x_1^B = 3$ and $E^B * x_2^B = 2$ for firm B.[11] Next, if we pool these adjusted inputs and outputs we have $y = 10$, $x_1^U = 5$ and $x_2^U = 5$. After adjusting and pooling the individual inputs, we again look at the amount by which we can proportionally reduce these inputs compared to the industry frontier. In this example, this delivers $E^{U^*} = 0.8$ as $E^{U^*} * x_1^U = 4$ and $E^{U^*} * x_2^U = 4$. This implies the following decomposition: $E^U = 0.4$, $E^{U^*} = 0.8$ and $T^U = 0.5$. Here we see that the relatively large savings that are due to E^U can be deceiving, and once individual inefficiency is controlled for the potential savings that are due to economies of scope are much smaller. The individual inefficiency scores $E^f = 0.5$ are reflected in the decomposition with $T^U = 0.5$.[12] The difference between a) and b) is that in a) the adjusted overall gains are zero, whereas in b) they are 20%, and this shows how the mechanics of generating synergies of this method works. Specifically, this demonstrates how firms can benefit from each other when they have comparative advantages in different inputs.

The asymmetric case c) is similar to case b) except that now firm B requires $x_1^B = 8$. Conducting the same exercise as in b) delivers $E^{U^*} = 0.8$. However, note that after multiplying the pooled and adjusted input $x_1^U = 6$ with $E^{U^*} = 0.8$ we obtain 4.8 which does not reach the frontier which is located at $x_1 = 4$. So there is further room for additional reductions of 0.8 in this input dimension. This further reduction is captured by the input slacks, which appear when the merging firms do reach the frontier in all input dimensions, after adjusting for individual inefficiency and proportionally reducing the inputs by E^{U^*}

Marginal cost reduction

As E_k^U represents the input specific synergy gains from merger U, we can reduce the output k of the merged entity by $(1 - E_k^U) * 100\%$. The new marginal cost of the merged entity U is given by:

$$\bar{mc}_{jt} = \begin{cases} mc_{jt} - \left(\sum_{k\in K} \left(1 - E_k^U\right) \hat{\gamma}_{f(j)c(j)k} g_{kt}\right) & for\ j \in U, \\ mc_{jt} \quad otherwise, \end{cases} \tag{1.24}$$

[11] E^f represents the individual inefficiency score. Here we have $E^A = E^B$ because of a symmetric example. However, in general, this is not the case.

[12] In this example, individual inefficiency scores E^f coincide with T^U. This is due to the constructed symmetric example and does not hold in general.

where $\hat{\gamma}_{f(j)c(j)k}$ is the estimated input k from the cost estimation that is used by firm f to produce output c. This estimated input is multiplied by the possible reduction $\left(1 - E_k^U\right)$, and the input price g_{kt}, as we want to subtract only the savings in Euros from the initial marginal costs mc_{jt}.

Price effects

In this section, we present the general methodology used to assess the effects of the change in the industry structure and cost savings on prices. The simulation is performed using the following program:

$$\min_{\left\{p_{jt}^*\right\}_{j=1,\ldots,J}} \left\| p_t^* - \Gamma\left(p_t^*, I_1^*, ..., I_{F-1}^*\right) - \gamma\left(p_t^*, I_1^*, ..., I_{F-1}^*\right) - \bar{mc}_t \right\|, \tag{1.25}$$

where $\Gamma\left(p_t^*, I_1^*, ..., I_{F-1}^*\right)$ and $\gamma\left(p_t^*, I_1^*, ..., I_{F-1}^*\right)$ are, respectively, the margins of manufacturers and retailers as a function of new equilibrium prices and new the ownership structure. The first three terms represent the marginal costs derived from the new price equilibrium, which should be equal to the new marginal costs $\bar{mc}_t$ in (24).

As we aim to isolate the price variation for the two opposite forces, we also simulate the price effects for two benchmark cases. On the one hand, we use the change in the industry structure only, then $\bar{mc}_{jt} = mc_{jt}$, and on the other hand, we use marginal cost savings only, then $\left(I_1^*, ..., I_{F-1}^*\right) = (I_1, ..., I_F)$.

We also compare our results to the common practice of reducing all marginal costs by 5% post merger. In this case, the new marginal costs post merger $\bar{mc}_{jt}$ are equal to $0.95 * mc_{jt}$ for merging firms and mc_{jt} for non-merging firms.

Furthermore, it is interesting to compare the impact on consumer surplus and the profit of the industry with and without synergies. Following Bonnet and Dubois (2010), consumer surplus is given by:

$$CS_t\left(p_t\right) = \frac{1}{|\alpha_i|} ln\left(\sum_{j=1}^{J} exp\left[V_{ijt}\left(p_t\right)\right]\right). \tag{1.26}$$

1.4 Empirical results

First, we discuss the results of the demand estimation, implied price-cost margins and the marginal costs. Then we present the results of the Data Envelopment

Analysis and the effects of the four proposed mergers.

1.4.1 Demand estimates, price elasticities and margins

We estimated a random coefficient logit model over a random subsample of 100,000 observations using a simulated maximum likelihood method, and these results are reported in Table 3. The price coefficients have the expected negative sign. We distinguished between national brands and private labels. The mean disutility of prices is higher for national brands with -3.05 and -2.33 for private labels, which means that consumers react more to price variations when they buy national brands than store brands. The standard deviation is estimated to be 0.93, such that only an infinitesimal part of α_i is positive. The error term of the price equation is positive and significant, meaning that the unobserved variables that explain the prices are positively correlated with the utility of buying a product. This is consistent with display and advertising costs, for example.[13]

The demand estimation allows us to compute own- and cross-price elasticities. Table 1.4 presents the average own-price elasticities across manufacturers and categories. Mean own-price elasticities range from -2.98 to -5.89. Private label products have the most elastic demand in all product categories. In the yoghurt category, manufacturers 1 and 3 have the least elastic demand. The second column suggests that no national brand producer has more market power than the others in the other dairy desserts category. In the third category, manufacturer 5 enjoys the most market power. Similar own-price elasticities were found by Draganska and Jain (2006) and Bonanno (2012) that ranged between -2.45, -6.25 and -1.4, -6.86, respectively.

The average implied price cost margins are reported in the Appendix in Table A3 across categories and manufacturers. On average, the total margin is 37% for yoghurt products, 32% for other dairy desserts and 29% for fromage frais and petit suisse products. Manufacturer 1 has the strongest market power given that the average margin of its products are the highest for the three categories, 49%, 40% and 44%, respectively. Manufacturer 3 exhibits the second largest margins with respect to the other firms. The lowest margins are for the private label products in each category. In total, the margins range between 17% and 49%. Other authors find similar margins: Bonanno (2012) recovers margins that range between 16% and 68%; and Villas-Boas (2007) finds margins that range between 12.8% and 45.8% in

[13]The estimation results of the price equation are in the Appendix Table A2. We chose milk, diesel oil and plastic prices as instrumental variables. They are all significant.

the preferred supply model. Regarding the results on marginal costs in the Appendix in Table A4, they seem to be consistent. In total, manufacturer 1 has the lowest marginal costs closely followed by manufacturer 3 and PL products. These three manufacturers have very similar low marginal costs for yoghurts and other dairy desserts. In the yogurt category, the marginal cost is higher for manufacturers 5 and 6 as they also use other milks that are more expensive than cow milk, such as soy milk or sheep and goat milks. As expected, the marginal costs of producing fromage frais and petit suisse and other dairy desserts are higher.

1.4.2 Cost function and DEA

For the estimation of the cost function, we use input prices from the French National Statistic Office (INSEE). We remain consistent with economic theory, as in Gasmi et al. (1992), we impose the positivity of the parameters δ_{fck}, and therefore, we use a non-linear least square method. We have chosen to use milk and salary as inputs, as they are the main inputs required to produce dairy products.

We present the results of the DEA in Table 1.5. In total, the DEA is performed with 13 initial firms that are shown in the upper part of Table 1.5 and 78 constructed mergers of which we show three selected mergers at the bottom part of Table 1.5. In total, we have 91 decision-making units with two inputs and maximal three outputs.

For the initial firms, we show the first stage efficiency scores E^f that are used to adjust the inputs before they are pooled to construct the mergers. A value of 1 means that this firm is part of the industry frontier, which is the case for five out of 13 firms.

At the bottom part in columns 2-4, the merger specific efficiency measures are given.[14] The first E^{U^*} in column 2 shows the linear efficiency scores from the DEA. For the mergers, this score is already adjusted for individual inefficiency. Thus, all inputs of a given firm that is part of the merger can be reduced by multiplying the inputs with the respective efficiency score and keeping the output constant. The implied input specific efficiency scores are given in columns 3 and 4 for milk and salary, respectively. These are also the scores that are used to calculate the savings in marginal costs in column 5. In the three mergers we consider, the input specific efficiency scores are equal to the linear efficiency score in column 2. This means that

[14] We use the bootstrap estimator presented in Kneip et al. (2008). In each bootstrap sample, we use 50% of the original sample size as the naïve approach leads to inconsistent results, as pointed out by Kneip et al. (2008)

there is no further input specific reduction in inputs possible.

The last column shows the average adjusted potential savings in marginal costs in percent, and the standard deviation across products, and time periods within the merged entity are shown in parenthesis. Even though the input savings are the same for a given product category and firm, the percentages vary as the original marginal costs are different. The savings of the selected mergers range between 1.88% and 7.84%.[15]

1.4.3 Merger simulation and welfare

We consider three potential mergers. The initial HHI in the French dairy dessert industry is 1,160.[16] The M1-M3 merger combines the pre-merger market shares of 24% and 8%, respectively. The change of the HHI is more than 250 and the initial HHI is above 1,000, such that this merger would be of concern for the Competition Authority according to the merger guidelines. This merger is expected to produce 7.84% savings in marginal costs. The M1-M11 combines pre-merger market shares of 24% and 4%, respectively, and is expected to produce savings of 1.88%. The change in HHI is 195. Manufacturers 8 and 9 both have 6% pre-merger market shares. Efficiency gains are expected to be 3.37% with a relatively low change in HHI of 67.

We simulate the new price equilibria and derive effects on prices, firm profits and consumer surplus for the three different mergers. For each case, we perform four different simulations. We derive the new equilibrium with the new industry structure only, second, with the new industry structure and the reduced marginal costs, and third, with the reduced marginal costs only. In this way, we aim to disentangle the opposing forces on prices. We compare the results to the usual approach of ad hoc savings of 5%.

[15]Table A8 in the Appendix summarizes mean savings of all 78 possible bilateral mergers. First note that only about 56% of the possible mergers are predicted to produce any efficiency savings at all. If we neglect the mergers that produce relatively small efficiencies and only consider mergers with at least 3% savings, then only about 35% of the mergers produce synergy gains. This means that about 44% of the mergers do not produce efficiency gains and are likely to harm consumers. Other authors find similar results. Our results are consistent with Gugler et al. (2003) who make use of a large panel data set to analyze merger effects of companies world-wide that occured during the 15 years prior to 2003. They find that approximately 29% of the mergers produced efficiency gains.

[16]Note for the calculation of the HHI and the changes post-merger, we assume that each private label manufacturer is treated as a single firm.

We also calculate the pass-through rate, market share variation and examine in detail the profit changes for each firm.

Note that the effects of mergers on the economy are highly dependent on the substitution pattern between the products of the merging firms. For example, if the products are close substitutes, we would expect that a lot of competitive pressure is taken away by merging, and thus firms can increase prices relatively more compared to a case of weak substitutes. On the other hand, we could have the case where two firms merge with relatively high market shares but the offered brands are only weak substitutes. In this case, a merger may have a low impact on the level of competition as both firms compete more with the firms outside the merger. This is also recognized by the Competition Authority in the merger guidelines.

Table 1.6 summarizes the results of the four simulations for the three merger cases. Columns 2-4 represent the three different mergers. For each of the mergers, we analyze the effects on prices and profits of the industry, and the prices and profits of the merged and outside firms separately.[17] As we are interested in the effects of mergers on consumers, we also present the changes in consumer surplus.

Benchmark

When we take into account the change of the industry structure only and the reduced marginal costs only, these results can be viewed as benchmark results. The former is a worst case and the latter a best case scenario from the consumers' perspective. We see that without cost savings, all figures behave as expected in all three mergers. Industry prices[18] and industry profits increase and consumer surplus decreases as expected. The merged firms benefit more than the average outside firm. Also price reactions are larger for the merged firm than the average industry price changes.

The other benchmark case — when we do not take into account the change in industry structure and look at the effects of the reduced marginal costs only — also show consistent results. As in this setup, there is only the downward pressure on prices, we see that all prices of the merged entity and the outside firms decrease in all cases. Note that we would expect that only the firms that experience the reduced

[17]Note that we cannot identify the fixed fees; in which case we are then only able to compute industry profit, that is, the sum of retail and manufacturer profit

[18]Industry price changes are computed as the weighted average (by market share) of price changes over all products in the industry.

cost would benefit in this new market equilibrium and the effect on total industry profit is therefore ambiguous. Indeed, we see that indeed only firms that have lower marginal costs increase their profits. The higher the marginal cost savings, the more they benefit. The firms with cost reductions now have a larger margin and thus have an incentive to decrease prices in order sell to more consumers. All the outside firms lose in the new equilibrium. As a reaction to the price cuts by the firms with lower marginal costs, outside firms cut prices as well and lose, as they do not enjoy any cost reductions to compensate for the lower prices. Industry profits increase in all three cases. Consumer surplus always increases, which is consistent with economic intuition, as reduced costs should always benefit the consumers without changes in the industry structure.

Net effect

In this section, we discuss the results of the scenario when we take into account both the change in industry structure and the reduction of marginal costs.

The M1-M3 merger produces the highest efficiency gains of the three mergers we consider, and also represents the largest increase in market concentration. Table A5 shows that the merging parties are also the main competitors of each other according to the substitution patterns. We see that industry prices actually decrease by 0.33%. Merged firms decrease prices to a larger extent by 1.81%. Here, the upward pressure on prices can be outweighed by the marginal cost reductions. Industry profits increase by almost 6%. This large increase is driven by the increase in the merging parties profits. Table A11 shows profit changes for each firm. Manufacturers 1 and 3 gain 17.39% and 19.10%, respectively. Outside firms lose between 1.20% and 1.83%. Interestingly, the post-merger strategy for both merging firms appears to be different. Table A9 shows that manufacturer 1 decreases and manufacturer 3 increases its prices post merger.[19] The merging parties seem to shift consumers from manufacturer 3 to manufacturer 1 by this pricing policy post merger. Table A3 shows that manufacturer 1 has the higher margin and also the higher market share pre-merger giving the merged entity incentives for this kind of strategy. Consumer surplus in this scenario increases by 0.92%, which is a large difference to the merger scenarios without efficiency gains and the ad hoc 5% rule, which predicts a decrease in consumer surplus of 1.1% and a very slight increase of 0.17%, respectively.

The M1-M11 merger results in a medium change of industry concentration of

[19] A negative pass-through rate means that prices have increased after a cost reduction.

more than 195 of the HHI and low savings of less than 2%. The reductions in marginal costs are just enough to compensate for the increase in concentration, which means that industry prices show minimal change, and rise only slightly by 0.06%. Merging firms increase prices by 0.37%. Industry profits increase by 1.63%. Merging firms benefit more than outside firms. Table A11 shows that manufacturers 1 and 11 gain 4.63% and 2%, respectively. Outside firms gain around 0.3% each. The strategy employed is similar as in the merger between manufacturers 1 and 3. As shown in Table A9, the extremely negative pass-through rate of manufacturer 11 indicates a relatively large price increase post merger. At the same time, manufacturer 1 decreases its prices slightly. Again, this suggests that the merging parties make consumers switch from manufacturer 11 to 1. Manufacturer 1 has a much higher margin and a higher market share. Consumer surplus decreases by 0.16%, making the merger anti-competitive. The 5% ad hoc rule delivers completely different results. It overstates the pro-competitive effects and predicts an increase of 0.66% in consumer surplus.

The merger between manufacturers 8 and 9 increases concentration by only 67 in HHI. The merger produces medium cost savings of 3.37%. We expect this merger to increase consumer surplus, as the downward pricing pressure is likely to outweigh the upward pricing pressure, as the increase in concentration is relatively small. Industry prices decrease by 0.14%. Merging firms decrease prices by 1.87%. Industry profits increase by 0.60%. The merging parties largely benefit from this merger as they increase profits by 12.83%. Outside firms lose 0.70% on average. Table A9 shows that both firms decrease prices post merger. As expected, consumer surplus increases post merger by 0.38%. The 5% ad hoc rule overstates the pro-competitive effects.

A notable detail of the presented results is that marginal cost savings and concentration effects seem to have an equal weight for the market outcome post merger for all effects, except for the profits of the merged entities. This can be seen in Table 1.6, where for most effects the net effect is approximately an average of the two benchmark cases. However, the changes in profits of the merged entities are almost only driven by cost savings. The benchmark results show that profit changes, when we only take into account synergy gains, are almost the same as profit changes, when we take into account both forces. Even in the case of M1-M3 with a relatively large increase in concentration, the changes in profits are almost only driven by synergy gains. This is an important result, as the main incentive to merge is the expected profit change. Our results suggest that managers are more likely to neglect the positive effect of concentration, and instead focus on possible synergy gains as the main driver for profit changes. Furthermore, post-merger pricing can be complex and may involve asymmetric pricing in order to encourage consumers to switch from one firm to the other.

1.5 Conclusion

Evaluating mergers is one of the key tasks of competition authorities. Empirical tools have become more sophisticated, taking into account strategic effects as well as vertical structure and contracts used within the vertical structure. However, one of the main aspects in mergers are the potential synergy gains that can arise post merger. So far, merger simulation models rely on ad hoc assumptions about cost reductions, or the models are used to find the required cost reductions to make a merger worthwhile. Other research aims at quantifying synergy effects due to fixed costs savings by reducing the product space.

In this article, we present an integrated approach to estimate efficiency gains and incorporate them into the merger simulation. We estimate a structural demand and supply model taking into account the vertical structure of the market. The contribution of this article is to use Data Envelopment Analysis in a next step, in order to derive synergy gains of potential mergers, and incorporate them into the structural model. We implement this methodology in the French dairy dessert market, and we simulate the impact of some bilateral mergers, taking into account the new ownership configuration and marginal cost savings. We find that only about 56% of the mergers produce any synergy gains, meaning that roughly half of the mergers do not produce any synergy gains and are therefore considered to be anti-competitive. The average marginal cost reductions are 2.55%, which implies that the ad hoc rule of 5% savings overstates the pro-competitive effects of mergers in the French dairy dessert industry. Depending on the marginal cost savings, the effects on prices, profits and consumer surplus differ. We find that the upward pressure on prices after a merger can be outweighed by the downward pressure, due to reduced marginal costs if savings are large enough. By isolating the concentration and efficiency effects on profits, prices and consumer welfare, we show that potential savings are just as important as market concentration. However, incentives to merge are fully driven by cost savings. Concentration effects on profits of the merging firms are relatively small compared to the effects that are due to cost savings. This suggests that policy makers require more reliant tools in order to screen mergers for their potential efficiency gains. Furthermore, our results suggests that firms may want to shift consumers from one firm to the other by asymmetric pricing strategies post merger.

Note that the predicted efficiency gains represent a maximal benchmark and should be regarded as potential cost savings. A natural next step for research would be to test the predictions and market outcome of our approach with real world mergers.

The limit of this analysis is that we cannot simulate long run effects of cost savings. As we have demonstrated, large savings are beneficial for consumers. However, this is only true in the short run. Large efficiency savings could prevent potential future market entry and may further reduce competition in the long run. Another limit of this analysis is that we do not account for product repositioning after a merger. As Mazzeo et al. (2014) find, merging firms may have incentives to reduce the number of products if the products are close substitutes. Another factor we do not take into account is the possibility of imposed remedies for the merging parties that are supposed to weaken the anti-competitive effects.

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1.6 Tables

Table 1.1: Market Shares by Producer and Category

	Category			
	Yoghurt	Dessert	Petit Suisse	Total
Manufacturer 1	0.38	0.21	0.13	0.24
Manufacturer 2	0.08		0.20	0.12
Manufacturer 3	0.11	0.22		0.08
Manufacturer 4	0.02	0.003	0.08	0.04
Manufacturer 5	0.02	0.05	0.03	0.03
Manufacturer 6	0.03			0.01
Manufacturer 7	0.07	0.09	0.11	0.09
Manufacturer 8	0.04	0.05	0.07	0.06
Manufacturer 9	0.05	0.07	0.06	0.06
Manufacturer 10	0.09	0.13	0.11	0.11
Manufacturer 11	0.03	0.04	0.05	0.04
Manufacturer 12	0.04	0.06	0.05	0.05
Manufacturer 13	0.05	0.08	0.09	0.08

Table 1.2: Average Retail Prices by Producer and Category

	Category		
	Yoghurts	Desserts	Petit Suisse
Manufacturer 1	1.90	2.44	2.00
	(0.59)	(0.26)	(0.12)
Manufacturer 2	2.19	.	3.02
	(0.43)	.	(0.51)
Manufacturer 3	1.75	2.20	.
	(0.15)	(0.19)	.
Manufacturer 4	2.45	2.25	4.42
	(0.27)	(0.04)	(0.39)
Manufacturer 5	3.30	2.83	6.28
	(0.39)	(0.29)	(0.69)
Manufacturer 6	2.92	.	.
	(0.16)	.	.
PL products	1.33	1.76	2.49
	(0.13)	(0.15)	(0.40)

Standard deviation are in parenthesis and represent variation across brands and periods.

Table 1.3: Demand Estimates

Parameter	Coefficient	StD
Price		
Average for NB products	-3.0511	0.0009
Average for PL products	-2.3376	0.0009
Standard deviation	0.9363	0.0001
Error term of the price equation	1.2820	0.0008
Brand fixed effects	Yes	
Retailer fixed effects	Yes	

Table 1.4: Own Price Elasticities

	Yoghurts	Other Dairy Desserts	Fromage Blanc and Petit Suisse
Manufacturer 1	-3.67 (0.48)	-4.12 (0.09)	-3.94 (0.10)
Manufacturer 2	-4.05 (0.17)		-4.10 (0.13)
Manufacturer 3	-3.67 (0.26)	-4.06 (0.12)	
Manufacturer 4	-4.14 (0.07)	-4.11 (0.03)	-3.59 (0.18)
Manufacturer 5	-4.05 (0.14)	-4.17 (0.05)	-2.98 (0.18)
Manufacturer 6	-4.18 (0.04)		
PL products	-4.19 (0.29)	-4.98 (0.24)	-5.89 (0.34)

Standard deviations are in parenthesis and represent variation across brands, retailers and periods.

Table 1.5: DEA - Constant Returns to Scale

Manufacturers	E^f	Manufacturers		E^f
Manufacturer 1	0.57	Manufacturer 8		0.92
Manufacturer 2	0.97	Manufacturer 9		0.75
Manufacturer 3	0.85	Manufacturer 10		1
Manufacturer 4	0.91	Manufacturer 11		0.78
Manufacturer 5	1	Manufacturer 12		0.77
Manufacturer 6	1	Manufacturer 13		1
Manufacturer 7	1			
Mergers	E^{U^*}	E^{Um}	E^{Us}	**Average MC Savings**
M1 - M3	0.86	0.86	0.86	7.84% (3.78)
M1 - M11	0.97	0.97	0.97	1.88% (0.97)
M8 - M9	0.92	0.92	0.92	3.37 % (2.03)

Standard deviation across products and periods in parenthesis.

Table 1.6: Merger Effects

Effects\Merger	Manufacturers M1 - M3	M1 - M11	M8 - M9
Cost Savings	7.84% (3.78)	1.88% (0.97)	3.37% (2.03)
HHI (ΔHHI)	1550 (390)	1355 (195)	1227 (67)
Prices			
Industry			
Only Merger	0.56% (0.22)	0.25% (0.01)	0.07% (<0.01)
Merger+Sav	-0.33% (0.03)	0.06% (0.01)	-0.14% (0.01)
Only Sav	-0.97% (0.02)	-0.20% (0.01)	-0.22% (0.01)
5 Percent	-0.04% (0.02)	-0.28% (0.01)	-0.22 (<0.01)
Merged firms			
Only Merger	3.08% (0.07)	1.58% (0.05)	0.95% (0.2)
Merger+Sav	-1.81% (0.18)	0.37% (0.04)	-1.87% (0.05)
Only Sav	-5.34% (0.15)	-1.26% (0.04)	-2.93% (0.07)
5 Percent	-0.22% (0.09)	-1.77% (0.06)	-2.90% (0.04)
Profits			
Industry			
Only Merger	1.51%	0.87%	0.33%
Merger+Sav	5.80%	1.63%	0.59%
Only Sav	3.82%	0.69%	0.17%
5 Percent	4.15%	3.06%	0.67%
Merged firms			
Only Merger	0.59%	0.39%	0.11%
Merger+Sav	17.80%	4.36%	12.83%
Sav	16.88%	3.93%	12.68%
5 Percent	11.27%	11.93%	19.59%
Outside firms			
Only Merger	2.08%	1.10%	0.36%
Merger+Sav	-1.73%	0.30%	-0.70%
Sav	-4.38%	-0.88%	-1.15%
5 Percent	-0.33%	-1.25%	-1.33%
Cons. Surplus			
Only Merger	-1.1%	-0.58%	-0.19%
Merger+Sav	0.92%	-0.16%	0.38%
Only Sav	2.40%	0.47%	0.62%
5 Percent	0.17%	0.66%	0.71%

Standard deviations across brands, retailers and periods are in parenthesis

1.7 Appendix

Table A1: Merger Gains Examples CRS

Entity	In and Outputs y	x_1	x_2	Linear Reduction E^f	E^U	E^{U^*}	T^U	Input Slacks S_1^U	S_2^U	E_1^U	E_2^U
Frontier	10	4	4								
	5	2	2								
a)											
Firm A	5	5	5	0.4							
Firm B	5	5	5	0.4							
Merger	10	10	10		0.4	1	0.4	0	0	1	1
b)											
Firm A	5	4	6	0.5							
Firm B	5	6	4	0.5							
Merger	10	10	10		0.4	0.8	0.5	0	0	0.8	0.8
c)											
Firm A	5	4	6	0.5							
Firm B	5	8	4	0.5							
Merger	10	12	10		0.4	0.8	0.5	0.8	0	0.66	0.8

Table A2: Estimates of the price equation

Parameter	Coefficient	Standard Error
Plastic price	0.011	0.005
Diesel oil price	0.006	0.002
Milk price	0.030	0.001
Brand fixed effects	Yes	
Retailer fixed effects	Yes	
F test of instrumental variables (p value)	1648.89	0.00

Table A3: Margins in Percent

Producer	Yoghurts	Other Dairy Desserts	Fromage blanc and Petit Suisse
Manufacturer 1	49.26 (9.58)	40.40 (2.30)	44.63 (1.86)
Manufacturer 2	25.32 (1.24)		25.04 (0.78)
Manufacturer 3	39.39 (2.22)	35.25 (1.78)	
Manufacturer 4	24.34 (0.42)	24.51 (0.20)	28.19 (1.42)
Manufacturer 5	24.81 (0.92)	24.12 (0.32)	33.90 (1.83)
Manufacturer 6	23.99 (0.22)		
PL products	24.19 (1.48)	20.15 (1.05)	17.16 (0.91)
Total	36.78 (12.90)	32.34 (8.04)	28.99 (8.94)

Standard deviation across brands, retailers and periods are in parenthesis.

Table A4: Marginal costs in Euros per liter

Producer	Yoghurts	Other Dairy Desserts	Fromage blanc and Petit Suisse
Manufacturer 1	1.02 (0.48)	1.46 (0.21)	1.11 (0.10)
Manufacturer 2	1.64 (0.34)		2.26 (0.36)
Manufacturer 3	1.06 (0.13)	1.43 (0.16)	
Manufacturer 4	1.86 (0.21)	1.70 (0.04)	3.17 (0.23)
Manufacturer 5	2.48 (0.27)	2.15 (0.22)	4.14 (0.36)
Manufacturer 6	2.22 (0.12)		
PL products	1.02 (0.12)	1.41 (0.14)	2.07 (0.36)
Total	1.36 (0.59)	1.55 (0.31)	2.45 (0.99)

Standard deviations across brands, retailers and periods are in parenthesis.

Table A5: Cross Price Elasticities NB

\	Man 1	Man 2	Man 3	Man 4	Man 5	Man 6	PL Man
Man 1		0.023	0.027	0.019	0.016	0.021	0.028
Man 2	0.004		0.004	0.004	0.004	0.004	0.003
Man 3	0.043	0.037		0.031	0.026	0.035	0.044
Man 4	0.002	0.003	0.002		0.003	0.003	0.002
Man 5	0.001	0.001	0.001	0.001		0.001	0.001
Man 6	0.001	0.001	0.001	0.002	0.002		0.001
PL Man	0.016	0.015	0.016	0.014	0.012	0.014	

Table A6: Cross Price Elasticities PL

\	Man 7	Man 8	Man 9	Man 10	Man 11	Man 12	Man 13	NB Man
Man 7		0.002	0.002	0.002	0.002	0.001	0.002	0.001
Man 8	0.016		0.016	0.016	0.016	0.015	0.016	0.015
Man 9	0.001	0.001		0.001	0.001	0.001	0.002	0.001
Man 10	0.005	0.005	0.005		0.005	0.004	0.005	0.005
Man 11	0.010	0.010	0.010	0.010		0.009	0.010	0.009
Man 12	0.072	0.072	0.072	0.071	0.072		0.072	0.068
Man 13	0.005	0.005	0.005	0.005	0.005	0.004		0.005
NB Man	0.014	0.014	0.013	0.013	0.014	0.012	0.014	

Table A7: Estimated Inputs

Category	Yoghurts		Dairy Desserts		FB and PS	
Producer	Milk	Salary	Milk	Salary	Milk	Salary
1	0.78	0.01	1.48	0.10	1.35	0.06
2	0.27	0.08			1.15	0.03
3	0.78	0.02	1.63	0.07		
4	1.46	0.05	1.00	0.04	0.34	0.06
5	0.83	0.01	0.71	0.03	1.12	0.15
6	0.45	0.03				
PL Producer						
7	0.03	0.01	0.50	0.01	0.50	0.07
8	0.52	0.04	0.48	0.01	0.68	0.04
9	0.71	0.05	0.37	<0.01	1.04	0.04
10	0.40	0.01	0.92	0.02	0.61	0.02
11	0.41	0.06	0.71	0.01	0.89	0.03
12	0.36	0.02	1.42	0.07	0.73	0.01
13	0.41	0.03	1.57	0.04	0.32	<0.01

Table A8: Summary Savings

	Fraction of Mergers	Av. MC Savings
All	100%	2.55%
Savings>0	56.41%	4.51%
Savings>1%	51.28%	4.90%
Savings> 3%	34.62%	6.16%
Savings> 5%	17.95%	8.27%

Table A9: Pass Through

Merger\	Scenario Merger+DEA	Only DEA	Merger+ 5%
M1 - M3 (7.84%)			
M1	0.35 (0.41)	0.95 (0.12)	0.36 (0.23)
M3	-0.36 (0.73)	1.09 (0.05)	-0.75 (0.20)
Average	0.11 (0.63)	1 (0.12)	-0.01 (0.57)
M1 - M11 (1.88%)			
M1	0.17 (0.49)	0.96 (0.12)	0.75 (0.14)
M11	-5.53 (2.99)	1.04 (0.01)	-0.68 (0.29)
Average	-0.14 (1.53)	0.97 (0.12)	0.67 (0.36)
M8 - M9 (3.37%)			
M8	0.53 (0.97)	1.02 (0.02)	0.73 (0.07)
M9	0.18 (0.24)	0.99 (0.08)	0.75 (0.08)
Average	0.35 (0.13)	1.01 (0.06)	0.74 (0.07)

Standard deviations across brands, retailers and periods are in parenthesis

Table A10: Market Share Variation

Merger\	Scenario Only Merger	Merger+DEA	Only DEA	Merger+ 5%
M1 - M3 (7.84%)				
Outside	2.15% (0.23)	-1.78% (0.20)	-4.55 (0.42)	-0.32 (0.07)
Inside	-9.12% (6.19)	6.93% (16.37)	19.37 (17.16)	0.14 (7.26)
Average	-3.35% (7.10)	2.48% (12.24)	7.13 (16.94)	-0.1 (5.08)
M1 - M11 (1.88%)				
Outside	1.11% (0.09)	0.29% (0.04)	-0.89% (0.07)	-1.26% (0.11)
Inside	-2.65% (5.05)	1.28% (6.34)	4.13% (3.61)	7.71% (5.95)
Average	-0.19% (3.47)	0.63% (3.76)	0.85% (3.20)	1.85% (5.52)
M8 - M9 (3.37%)				
Outside	0.33% (0.04)	-0.66 (0.08)	-1.07 (0.13)	-1.20 (0.16)
Inside	-4.04% (0.17)	6.90 (6.37)	11.92 (6.59)	14.16 (5.12)
Average	0.17% (0.83)	-0.38 (1.88)	-0.59 (2.76)	-0.63 (3.07)

Standard deviations across brands, retailers and periods are in parenthesis

Table A11: Profits by Firm

Manufacturers\Merger	M1 - M3	M1 - M11	M8 - M9
Cost Savings	7.84% (0.36)	1.88% (0.18)	1.62 % (1.3)
Manufacturer 1	17.39%	4.63%	-0.71%
Manufacturer 2	-1.72%	0.30%	-0.69%
Manufacturer 3	19.10%	0.31%	-0.74%
Manufacturer 4	-1.39%	0.27%	-0.58%
Manufacturer 5	-1.20%	0.25%	-0.53%
Manufacturer 6	-1.78%	0.30%	-0.70%
Manufacturer 7	-1.81%	0.30%	-0.73%
Manufacturer 8	-1.83%	0.30%	12.86%
Manufacturer 9	-1.82%	0.30%	12.80%
Manufacturer 10	-1.79%	0.30%	-0.72%
Manufacturer 11	-1.83%	2.00%	-0.74%
Manufacturer 12	-1.79%	0.30%	-0.72%
Manufacturer 13	-1.79%	0.30%	-0.73%

Chapter 2

Local Market Structure and Consumer Prices: Evidence from a Retail Merger

Co-authored with Dennis Rickert and Joel Stiebale

2.1 Introduction

There has been a substantial rise in industry concentration and market power in most sectors in the US and Europe over recent decades (e.g., Autor et al., 2017; De Loecker and Eeckhout, 2017). This development is at least partly a result of merger and acquisition (M&A) activity which has gained importance in recent years with a combined worldwide deal value that exceeded $4 trillion for the first time in 2015.[1] Retail industries have been particularly affected by this development. For instance, the four largest firms within US retailing sectors have increased the average share of sales by more than 15 percentage points within 20 years (Autor et al., 2017). In Germany, France, and the UK, the combined market share of the five biggest retailers exceeds 70% (Inderst, 2013), whereas in Northern European countries the three largest retailers hold combined market shares of around 90% (Allain et al., 2017). Since retail purchases constitute a high share of consumption expenditures, increasing retail concentration is a central topic both in economic policy and within public debates.[2] In this paper, we analyze the effects of changes in local market structure—induced by a merger between two German retail chains—on consumer prices.

Economic theory yields contrary predictions regarding the impact of M&As. On the one hand, mergers might allow firms to reduce competition in downstream markets and increase prices at the expense of consumers (von Ungern-Sternberg, 1996; Dobson and Waterson, 1997). On the other hand, a merger can can lead to efficiency gains which benefit consumers in the form of lower prices or new and improved products (Williamson, 1968). The net effects are particularly difficult to predict in the presence of buyer power in vertical relations which play an important role for retail markets. Specifically, merging firms might benefit from lower wholesale prices due to enhanced bargaining power vis-a-vis their suppliers (e.g., Galbraith, 1954; Chipty and Snyder, 1999) or size-related discounts (e.g. Katz, 1987; Scheffman and Spiller, 1992). Whether consumers benefit from efficiency gains or enhanced buyer power of merging firms, however, depends on suppliers' bargaining position

[1] See, for instance, `http://www.wsj.com/articles/2015-becomes-the-biggest-m-a-year-ever-1449187101`, accessed November 2, 2017. See Grullon et al. (2017) for a recent study that relates variation in concentration levels across industries and time to M&As.

[2] See, for instance, `https://www.economist.com/news/finance-and-economics/21727893-digital-age-protecting-customers-interests-harder-ever-market`, accessed November 2, 2017.

relative to other retailers (Inderst and Valletti, 2011) as well as characteristics of market structure and demand which determine pass-through of cost savings (Bulow and Pfleiderer, 1983; Weyl and Fabinger, 2013; Gaudin, 2016).

Due to mixed theoretical predictions, and the difficulty to match observed price patterns with counterfactual merger simulations, researchers have recently argued that more evidence from ex-post merger analysis is clearly needed (e.g., Angrist and Pischke, 2010).[3] Subsequently, a growing empirical literature has estimated the effects of M&As on prices and other outcomes and has produced mixed results. For instance, Hosken et al. (forthcoming) analyze 14 retail mergers and conclude that prices were raised after some mergers while they decreased or remained unchanged in other cases. Allain et al. (2017) report a significant price increase after a merger among French retailers, while Argentesi et al. (2016) find adjustments in the product portfolio after a merger in the Dutch retail market but no significant price changes within products.[4] The ambiguity of theoretical and empirical results is also reflected in heterogeneity of national competition authorities' decisions on mergers. For instance, both the European Commission and the Federal Trade Commission approved almost 90% of all proposed retail mergers without constraints (Allain et al., 2017). In contrast, based on concerns about consumer surplus and producer profits, Germany's federal cartel office (Bundeskartellamt) either denied or imposed strict remedies on all mergers that were recently proposed.[5]

The aim of this paper is to investigate the price effects of a merger between two German retail companies, where a large supermarket chain acquired a soft discounter. This merger was challenged by competition authorities but eventually cleared subject to remedies at the end of 2008.[6] Following the previous literature on ex-post merger evaluations, we start by estimating the average causal effect of the merger on consumer prices. We develop a unique identification strategy to dis-

[3]While there is a prominent and controversial debate about the usefulness of structural models versus quasi-experimental analysis in industrial organization in general and merger analysis in particular (e.g., Angrist and Pischke, 2010; Einav and Levin, 2010; Nevo and Whinston, 2010), most scholars agree that additional insights from retrospective merger analysis is important.

[4]See also the overview of related literature on mergers in various industries in Ashenfelter et al. (2014). Recent contributions that compare the results of merger simulations and retrospective analysis include Björnerstedt and Verboven (2016) and Friberg and Romahn (2015).

[5]See, for instance, Bundeskartellamt (2014a) on denied mergers or Bundeskartellamt (2005, 2010) on conditionally excepted mergers.

[6]Due to contractual agreements with the data provider, Gesellschaft für Konsumforschung (GfK, Germany), we are unable to name the merging firms, although we are allowed to state identifying information such as firms characteristics and year of the merger.

entangle price growth due to increasing market concentration from price declines that stem from cost savings. Further, we contribute to the literature by analyzing heterogeneous effects which can potentially explain the inconclusiveness of previous empirical studies. In particular, we investigate differences in price responses between discounters and supermarkets, between private label products and national brands, and how these responses vary with predicted changes in regional retail market concentration. By relating the direction of price effects to local market conditions, our study provides a potential explanation for the ambiguous results of previous ex-post merger evaluations.

We use a rich consumer-level panel data set which allows us to construct measures of prices per product and retailer for several thousand geographical markets at the municipality level and to control for a various aspects of regional heterogeneity. The empirical analysis focuses on 4 product categories: milk, yogurt, coffee, and toilet paper and thus includes a mix of differentiated and rather homogeneous products. There are several interesting features of the merger and German retail markets which we exploit in our identification strategy. First, the merger has been decided at the national level and was therefore unlikely to be related to regional pre-merger market characteristics. Further, most German retailers employ a regional pricing strategy that is adapted to local market conditions while decisions about warehousing and bargaining with suppliers are usually made at a higher geographic level. Finally, there are several regional markets with pre-merger presence of both acquirer and target, only one of the merging firms, and none of both firms, respectively. We use difference-in-differences (DiD) estimators to exploit geographical and cross-time variation of consumer prices within products sold by each retail chain.

Our baseline specification to identify price effects of the merger follows the previous empirical literature on retrospective merger analysis (e.g., Dafny et al., 2012; Allain et al., 2017; Ashenfelter et al., 2015; Argentesi et al., 2016). It is based on a comparison of geographical markets in which both acquirer and target operated in the pre-merger period to control markets that did not experience a change in market concentration. In a subsequent setup, we explore price changes through concentration effects in more detail. For this purpose, we compare local markets in which both firms were active pre-merger to a comparison group of geographical markets in which exactly one of the merging parties was operating. The former group of regions was affected both by a change in market concentration and by potential efficiency gains, while the latter group was presumably affected by nationwide cost savings in one of the retail chains but did not face systematic changes in local market concentration.

To investigate the importance of price declines due to cost savings, we compare markets in which one but not both merging firms operated in the pre-merger period to markets in which neither acquirer nor target were active. The former but not the latter group was potentially affected by nationwide cost savings of the merging firms and corresponding price responses of insiders and outsiders. Since none of both groups experienced a change in local market concentration, price effects identified by our DiD estimator in this setup are likely due to changes in costs that are partly passed on to consumers. These cost savings may stem from economies of scale and scope, e.g. from sharing warehouse capacities, or from higher bargaining power vis-à-vis upstream manufacturers.

Our results indicate that merging retailers and their competitors raised average consumer prices in affected markets. While the estimated price changes are on average quite small (about 0.4%), prices increased substantially in regions with high predicted changes in market concentration. For the region with the largest increase in retail market concentration in our sample, the estimated price increase amounts to 7.04% for supermarkets which is substantial given that estimates of recent retail markups lie below 0.3 for the median firm (see, for instance, Hottman, 2017). Price adjustments are concentrated in products sold by supermarkets, but there is little evidence for changes in pricing strategies of discounters. A possible explanation for this heterogeneity is the large overlap between the acquired soft discounter's product portfolio and those of supermarkets. Further, German hard discounters are likely to enjoy considerable market power in many product categories. We also investigate the effects of remedies imposed by the German cartel office which involved the sale of target firm's retail stores to a competitor in regional markets with high pre-merger market shares of acquirer and target. The results suggest that prices were not less likely to rise in these markets, indicating that imposed remedies might not have been sufficient to prevent anti-competitive effects of the merger.

Comparing price effects of the merger across alternative treatment and control groups, we find—as expected—significant downward pressure on prices in markets that are likely to be affected by potential efficiency gains. The negative effect on prices is higher in initially more competitive markets which is consistent with imperfect pass-through of cost savings. In contrast, firms raised prices in regions with pre-merger overlap significantly, and this effect increases with expected changes in local retail market concentration. Consistent with price increases due to market power, estimated relative price growth is higher when we compare regions with changes in market concentration to regions that experienced no change in market structure but were likely to be affected by national-wide cost savings to a similar extent as the treatment group. Our results are robust to controlling for a large set of potentially confounding variables including retailer-product-region fixed effects, retailer-year fixed effects, brand-year fixed effects and various factors capturing variation in demand across geographical markets and time. We also obtain similar results when we use alternative regional market definitions, exclude control markets that are geographically close to treatment markets, or use a propensity score reweighting estimator.

The rest of this paper is organized as follows. Section 2 discusses the German retail market and the merger case. In section 3, we provide a description of our consumer-level panel data set, section 4 details our identification strategy. Results of the empirical analysis are presented in section 5, and section 6 concludes.

2.2 The merger and the German retail market

In this section, we first provide detailed background information on the merger in section 2.2.1 before we describe pre- and post-merger market structure (section 2.2.2) and characterize the local component of price competition among retailers (section 2.2.3).

2.2.1 The merger and the German retail market

The German retail sector has developed to a highly concentrated market structure over time. Induced by an expansive M&A strategy, the five largest retailers in Germany have increased their market shares in the two preceding decades from 50% to over 80% in 2014 (Inderst, 2013), which is above the average of 70% in other Western European countries and well above the US average of 33% (Allain et al., 2017). Prominent examples for recent mergers in Germany are the cases of Edeka/Trinkgut (Bundeskartellamt, 2010), Edeka/Tengelmann (Bundeskartellamt, 2014a), and Wasgau/Rewe.[7] We analyze the merger of two retailers $R1$ and $R2$ with pre-merger market shares of 25% and 5% which was proposed at the end of 2007

[7] `http://www.bundeskartellamt.de/SharedDocs/Meldung/DE/Pressemitteilungen/2013/29_04_2013_Rewe_Wasgau.html`

and was approved by competition authorities in midyear 2008.

The three largest competitors, outsiders O1, O2 and O3 have market shares of 20%, 15% and 15%, respectively. The acquirer $R1$ is a multi-line retailer with two different retailing formats, supermarkets and discounters, which we label $R1^S$ and $R1^D$. The target $R2$ can be classified as a soft discounter which charges relatively low prices compared to supermarkets but has a broader assortment and sells a relatively high share of national brands compared to hard discounters. Since aggregated post-merger market shares exceeded the safe-harbor threshold of 22% (Competition Commission, 2008), the merger was in the focus of the national cartel office which identified regions where the firms had large market shares and concluded that competition would be potentially distorted by the merger. The merger was approved under the condition that $R1$ divests and sells 378—out of 2700 stores in question—to outsider $O1$. $R1$ converted and relabeled 1800 stores to $R1^D$, the remaining stores kept their former label $R2$ but were effectively under the control of $R1$. The outsider $O1$ pursued the same strategy in remedy regions, where it acquired the target's stores. It relabeled the acquired stores into its own soft discount retail chain $O1^D$. Basically, the imposed remedies implied a second merger between $O1$ and $R2$ in remedy regions. Consequently, we will treat remedy regions as part of the treatment group in our baseline specification. However, in an extension of our analysis, we also investigate remedy and non-remedy regions separately.

2.2.2 Pre- and post-merger market structure

34 different retailers are active on the German grocery retailing market which can be grouped into three formats: Discounters, drugstores, and supermarkets with market shares of 51.76%, 3.02%, and 44.56%, respectively. Furthermore, there are some specialized retailers, such as, for instance, cash-and-carry stores, pharmacies, and online retailers. However, neither these specialized retailers, nor internet purchasing can be regarded as close substitutes for grocery purchases at supermarkets and discounters during our sample period. Thus, we exclude all products from these distribution channels from our analysis. We define insiders as the two merging firms, which are a supermarket and a discounter and refer to the remaining firms as outsiders.

Table 2.1 shows average market shares per category distinguishing between national brands and private labels, and retailing format pre and post merger. The table shows that private labels are an important element of the market. Particularly, discounters' assortment consist of a high share of private labels, but we see that supermarkets also offer a high proportion of private labels in some products categories. Discounters have the highest market shares in markets for toilet paper, yoghurt and milk where they mostly sell their private label products. Supermarkets dominate in the coffee market by selling national brands.

Table 2.2 shows average prices per category separately for treatment and control group and retailing format before and after the merger. We see that pre and post merger, in treatment and control group, discounters have lower prices in all categories. Post-merger, supermarkets slightly raise their prices for toilet paper, yogurt and milk in treatment markets relative to control markets.

2.2.3 Local market definition and national bargaining

In order to identify the causal effect of the merger on retail prices, an accurate market definition is essential—both with respect to retail sales to consumers as well as regarding procurement from suppliers. We argue that similar to the US (Dafny et al., 2012), the Netherlands and France (Allain et al., 2017)—but unlike in the UK (Dobson and Waterson, 2005)— German retailers adopt a local pricing strategy (see also Bundeskartellamt, 2014a) and that retail purchasing takes place on national procurement markets. This assumption is supported by anecdotal evidence from two large German retailers—Edeka and Rewe. Edeka and Rewe evolved from former buying cooperatives of local merchants, which were subsequently transformed into national retail chains with centralized headquarters. Due to this historical development, local merchants can independently set prices and choose assortment, while national retailers bundle purchasing activities of local retailers.

These characteristics are also in line with information from one of the acquiring firm's web page. $R1$ owns 11,400 stores operated by roughly 4,500 independent merchants who adapt the day-to-day business activities to local markets.[8] The retail brand $R1$ has evolved from regional cooperatives that were created for the purpose of joint purchasing activities. At the time of the merger, there have been seven regional wholesale cooperatives—formed by previous regional buying groups—, who deliver items to the stores of independent merchants and coordinate central issues regarding distribution and sales. These wholesale cooperatives may also own retail outlet stores and production facilities. The distribution is effected from 38 dis-

[8]Due to confidentiality agreements, we are neither allowed to display names nor links to homepages of the merging parties.

tribution warehouses managed by the seven wholesale cooperatives. Furthermore, there is a central headquarter coordinating commodities transactions ("Nationales Warengeschäft") at the national level. The headquarter employs a national purchasing strategy for many food and all non-food product categories.

Two main conclusions can be drawn from the facts discussed above which play an important role in our identification strategy described in next section 4. First, local merchants receive their stock from central distribution warehouses which are coordinated by the national headquarters. Wholesale prices are thus likely to be determined at the national level if there is no price discrimination among local merchants. The assumption is in line with practices of the German antitrust authorities who define procurement markets at the national level (Bundeskartellamt, 2014b, p.132). Another indication for national procurement markets are so-called "wedding rebates", where some retailers were under suspicion of demanding better purchasing conditions for all stores after the acquisition of the target which were debated extensively in public.[9] Second, despite national wholesale prices, local retailers can set prices, including regular prices and discounts, independently. Hence, local retail prices are likely to vary across regions within retailers even for products with a wholesale list price that is determined at the national level.

Since local market definition (and the nature of retail price setting) is crucial to merger evaluation, we provide descriptive insights on the local dimension of retail price setting practices from our micro data. For this purpose, we regress prices on retailer-brand fixed effects for each product category and time period. This simple regression yields an R^2 indicating the explanatory power of the national pricing component at the retail-brand level. The remaining residual variation of the regression is the share of variance which cannot be explained by national pricing strategies. Figure 2.1 plots $1 - R^2$ over time and product categories indicating the decomposition of price variation into national and regional variation. For all product categories, we find a high share of variance which can be explained by local components, ranging from roughly 25% for milk and yogurt to around 65% for toilet paper. To provide evidence on the determinants of price differences across regions, we also regress prices on regional characteristics. Table B1 shows that regional prices vary with local market conditions such as the share of households with children, population density, and market concentration. While this regression

[9]See for instance `http://www.bundeskartellamt.de/SharedDocs/Meldung/EN/Pressemitteilungen/2013/24_07_2013_Edeka.html?nn=3591568`.

does not allow to infer causal relationships between prices and market structure variables, it indicates that geographic variation in retail prices is related to local demand conditions.

The Bundeskartellamt ackknowledges these local components of retail competition and defines 345 local catchment areas, which are the basis for decisions on mergers or abuse of dominance (Bundeskartellamt, 2007). In contrast to this rather broad market definition, we define local markets at a more disaggregated level which corresponds to the classification of municipalities. According to this classification, there are roughly 12,000 local markets and the set of competitors therefore contains all stores located in this municipality. Municipalities have an average size of 32.10 square kilometers (median of 18.10 km^2) and a standard deviation of 40.58 km^2.[10] Our definition of local markets is also relatively narrow compared to the definition of the European competition commission (European Commission, 1999) who proposes to define retail market using circles around stores with radii that correspond to 20 minutes driving time by car (roughly 15-20 kilometers) since we believe that most German consumers are likely to do their shopping within a smaller neighborhood around their residence. However, our market definition could be too narrow for some rural areas but at the same time be too wide for big cities such as Hamburg or Berlin. Therefore, we conduct two sensitivity analyses with respect to market size definition. First, we exclude all regions within a 15km circle around the treatment group. As these regions are potentially affected by the observed mergers, they should not be contained in the control group. Second, we exclude all urban regions from the estimation sample. If regions are defined too widely, we erroneously assign regions not affected by the merger to the treatment group, and our estimates capture a lower bound. Results presented in section 5.3 are consisted with this argument.

Based on these stylized facts discussed above, we develop an empirical strategy which is consistent with observed market characteristics and evidence on retailers' strategies.

[10]We have calculated these numbers using additional information received from DESTATIS `https://www.destatis.de/DE/ZahlenFakten/LaenderRegionen/Regionales/Gemeindeverzeichnis/Gemeindeverzeichnis.html`, accessed March 31, 2017.

2.3 Data

To estimate the effects of the merger, we exploit a rich data set consisting of a household panel survey complemented with regional information from additional sources. We present and describe this dataset in section 4.2 before reporting details on construction of products and prices in section 2.3.2.

2.3.1 Data description

The primary dataset is a representative survey of households distributed across all regions of Germany obtained by GfK Panel Services. The GfK Panel Services collect information on all transactions of up to 20,000 households which are selected to be representative of the German population with respect to geographical, social, and economic characteristics. This rich dataset entails two distinct features which makes it well-suited for the purpose of our analysis. First, all panel members track their entire purchase history using home-scanning devices. Thus, it contains detailed information on the name of the brand, the label type (national brand or private label), the retailer (e.g., supermarket, discounter, drugstore, or specialized shop), and type of product (including package size among other characteristics) as well as the actual transaction price (including any discounts and promotions). Thus, it gives a more accurate picture on household shopping behavior compared to checkout scanner data—which can only track purchases within a particular store—or datasets from other marketing agencies in Germany, which do not provide information on all discounters (see e.g., Draganska et al., 2011). Second, the data encompass detailed information on panel member characteristics, including the postal code of their residence, their yearly income, the number of children, and the job occupation.

We merge our data with information obtained from INKAR providing regional information at the level of counties ("Kreise") and municipalities ("Gemeinden"). For this purpose, we match postal codes with a municipality identifier for which we use a matching key provided by Deutsche Post, Germany's largest postal service company. Within our sample period, there have been a number of reforms, where postal codes were reallocated to other municipalities or where new municipalities have been created by merging existing ones. To address this issue, we retrospectively allocate postal codes to the definition of the year of 2010. Furthermore, there are some cases where the local postal code may belong to two (or more) municipalities. Since these postal codes are unlikely to be systematically related to the assignment of treatment and control group, we drop these cases, which leaves us with 78% of the total observations.

Having matched the postal code to the municipality level also allows to identify all purchases and prices within a regional market. While we lack precise information about the location of stores, we can assign consumer level purchases to regions and retail chains. We assume that a retail chain owns at least one store located in a specific region if we observe purchases at the retail chain in the region. Finally, we use additional data from the German cartel office—available at the 2-digit postal code level—to identify regions with remedies.

2.3.2 Construction of products and prices

Our empirical analysis focuses on four product categories: milk, yogurt, coffee, and toilet paper. The selected products reflect a mix of rather homogeneous base products: milk and toilet paper, more differentiated products: yogurt and coffee.

Observing a—albeit representative—subsample of the German population has the disadvantage that for some regions and some products we observe low (or sometimes zero) frequencies per day within a region. Consequently, to ensure a sufficient number of observations, we define products as the category-brand combination and aggregate the data to quarterly time periods. Between 2005 and 2010, we observe 3,019,952 purchases. The distribution of purchases across product categories is: 9.1% toilet paper, 9.2% coffee, 44.7% yogurt, and 36% milk purchases. On average, we observe 65 purchases per period and local market ranging between 1 and 9185 with a standard deviation of 224. To account for panel attrition—that is entry and exit of panel members into the sample—we restrict our observations to purchases of households who have been active at the beginning and at the end of the sample period.

We construct mean prices per product, retailer and region in Eurocents per unit of size. This unit of size depends on the product category. It is either grams or milliliters for food products, i.e., coffee and dairy products. For toilet paper, a per-unit price is used. Our price definition is the transaction price, which is the effective price paid at the checkout counter. In our baseline setting we build separate mean prices for target and acquirer. To check whether changes in prices reflect composition effects (e.g. because the merger target had fewer national brands in its portfolio than the acquirer), we treat the merging parties as a single firm in markets with pre-merger overlap as a robustness check which we discuss in section 2.5.3. As a further robustness check, we use prices weighted by the number of purchases as in Allain et al. (2017).

2.4 Empirical Strategy

The aim of this study is (i) the ex-post evaluation of price effects of the merger between retailers $R1$ and $R2$ and (ii) the decomposition of the overall effect into price changes due to changes in (a) local competition and (b) cost savings. For this purpose, we develop a novel and unique identification strategy which we discuss in section 2.4.1. Section 4.1.1 summarizes the basic assumptions of our identification strategy. Sections 4.1.2 and 4.1.3 describe the definition of our treatment and control groups for different market structures in more detail. First, we use a standard definition of treatment and control group in the retrospective retail merger evaluation literature defining the treatment group as any local market that experienced a change in local concentration after the merger. This approach identifies an average treatment effect assuming retailers do not employ national strategies (section 4.1.2). Since retailers' strategies might be neither completely local nor completely national, we adopt the framework of Allain et al. (2017) to measure the price effect which allows for merger effects at the national level in section 4.1.2. Within this framework, we are also able to analyze efficiency gains and market power effects by identifying regions where we expect market power to be a dominant force and regions where we expect efficiency gains to play an important role. Section 2.4.2 presents the empirical specification.

2.4.1 Identification

A simple before-after analysis is not sufficient to estimate the effects of mergers on prices. Observed price changes might also be attributed to shifts in demand or costs. Therefore, we aim to compare price changes around the merger to a counterfactual scenario in which no merger took place. For this purpose, we exploit the fact that neither the target nor the acquirer owned retail stores in each local market before the merger. Therefore, we compare regions which experienced a change in market concentration, i.e., markets in which both $R1$ and $R2$ were active before the merger, to markets without pre-merger overlap. Our identification strategy thus relies on the assumption that firms use a regional pricing strategy such that changes in product prices are independent across regional markets— conditional on a range of observable characteristics and fixed effects. The stylized facts in section 2.2.3 and the market definition exercises of the European competition commission (European Commission, 1999) as well as the German cartel office (Bundeskartellamt, 2014b) provide strong evidence for local pricing decisions, such that we can rule out pure national pricing strategies.

The causal price effect due to the merger is identified by the implementation of a simple difference-in-difference (DiD) estimator. The DiD approach compares pre- and post-merger prices of treated regions (i.e., regions affected by the merger) to pre- and post-merger prices in a control group. Taking double differences isolates the merger effect from other factors that might impact prices, such as characteristics of regional markets and retail-chain specific demand and cost shocks. However, the estimator relies on a parallel trend assumption, which implies that prices in treatment and control group would have moved identically in the absence of the merger. For this reason, we describe how we define local markets and their assignment to treatment or control group in the following subsections.

Treatment- and control groups in baseline scenario

Local markets are defined by the borders of municipalities. The dataset—albeit containing rich information about product and consumer characteristics—lacks information on the location of retail stores. However, it provides the location of every consumer at the postal code level, which allows to infer store locations by assuming that households buy products in the local market of their residence. We match the postal code to regional data at municipality level via a matching key. Subsequently, we are able to identify all purchases and prices within a local market and for each quarter. From purchases within the municipality we infer the location of insider firms $R1$ and $R2$. The set of competitors therefore contains all stores located in this municipality.

We follow the literature on retail merger evaluation (e.g. Houde, 2012) to define the **treatment group** as local markets affected by the merger. More precisely, the baseline specification of our treatment group contains regions which experienced a change in market concentration, i.e. markets in which both $R1$ and $R2$ were active before the merger. As the merger was approved under the condition to sell roughly 300 stores to competitors in some local markets, our definition of treatment groups includes local markets where the competitor acquired the target. Consequently, we add remedy regions to the treatment group, which implies that merging firms in these regions are referred to as $R1$ and $R2$.

Treatment groups are compared to the **control group** of markets without pre-merger overlap, where we assume parallel price trends for treatment and control group absent the merger. The broadest definition of a control group would contain all local markets without a pre-merger market overlap of acquirer and target. However, we follow Allain et al. (2017) and exclude regional markets that are located geographically close to treatment markets, which could be indirectly affected by the merger and thus contaminate results. To this extent, we assume that each store is located at the municipality center and define so-called catchment areas around these stores.[11] As a robustness check, we also construct an algorithm which (i) identifies the center of each municipality, (ii) calculates the distance to all other municipalities, and (iii) eliminates all municipalities within a distance of less than 15 kilometers. Results, however, are robust to this robustness check.

Figure 2(a) illustrates how we define treatment and control groups for our baseline specification. Treatment regions, which consist of local markets where both $R1$ and $R2$ were active, are colored in red. Control groups—colored in green—are defined as all products in markets without pre-merger overlap. According to this definition, control groups are defined as local markets with (i) only outsiders, (ii) outsiders and either $R1$ *or* $R2$, or (iii) either $R1$ *or* $R2$, but no outsiders. The circle around the treatment groups containing the arrow indicates the aforementioned robustness check, where we exclude counties which have no pre-merger overlap between $R1$ and $R2$ but share a border with a merger county or are located less than 15 kilometers away from a market directly affected by the merger (grey-colored local markets), to ensure that our control group is not contaminated by indirect merger effects (see section 2.5.3).

The definition of treatment and control group defined above, identifies price effects due to changes in local market power and regional efficiency gains—if there are any. This specification also serves as a simple test for this hypothesis. In case of pure national pricing, we expect to find no significant price differences between treatment and control group. However, although stylized facts and reduced-form

[11] We use the user-written STATA command opencagegeo (available from http://fmwww.bc.edu/RePEc/bocode/o/opencagegeo.pdf) to obtain longitude and latitude of each municipality.

regression provide strong evidence against a pure national pricing strategy, prices might be, at least partially, determined at the national level, and the internalization of competition effects may be adapted globally. If this was the case, insider firms potentially internalize competition externalities and efficiency gains regardless of their location implying that prices uniformly change at a national level. In this case, control markets in which at least one of the insiders operates are confounded. The subsequent section 2.4.1 provides a setup to address this effect.

Concentration effects, efficiency gains, and net effects with partly national strategies

The above specification incorporates insiders in some of the regional markets. If retailers' prices are affected by national retail strategies or efficiency gains, all local markets with stores that belong to acquirer's or target's retail chain may be affected by the merger. In this section, we first illustrate how we address this issue. The baseline scenario merely sums up efficiency gains and concentration effects, where the sign of the overall effect indicates which channel dominates. However, it does not allow disentangling the magnitude of price changes due to efficiency gains from those that stem from a change in market power. Therefore, we also present a novel identification strategy, which is based on modified treatment and control groups, to explore the effect of (the pass-through of) efficiency gains and concentration effects.

Insiders may have benefited from global efficiency gains, which are potentially (partly) passed-on to consumers. In order to learn more about the magnitude of the pass-through of global efficiency gains, we exclude all local markets from the control group which contain either $R1$ or $R2$ (see Figure 1(b)). Intuitively, we now compare markets with pre-merger overlap to a control group of unaffected markets without insiders. In this setup, the price effect is only identified by price adjustments of competitors since we estimate price effects at the market-retailer-brand level and insiders are not present in the control group. Thus, we compare prices of outsiders in treatment and control group. We expect to see price effects if outsiders respond strategically to the merger. Excluding regions with insiders from the control group eliminates the effect of potential regional efficiency gains and allows to focus on efficiency gains at the national level. Abstracting from regional efficiency gains—such as increase in local bargaining power or cost savings from local transportation—seems reasonable for the merger as we have outlined in section 2.2.3. Since $R1$ pursues a national purchasing strategy, distribution cost savings are likely to be internalized at the national level. The same applies to cost reductions that can be achieved through increased bargaining power with respect to suppliers. Even though the specification excludes both insiders, control markets are a reasonable comparison for treatment markets since (i) we have a sufficient number of cases (see table 2.3) and (ii) outsider $O1$—a close competitor with a similar business strategy and format as $R1$—is active in many of those regions. Further figures 2.3–2.6 show that trends in pre-merger prices are very similar between treatment and control markets. In this specification, price effects should be lower—compared to the baseline scenario—since insiders potentially benefit from nation-wide efficiency gains and these might be (partly) passed on to consumers. Such global cost savings are likely to be partially canceled out in the previous specification, whereas they are likely to affect only the treatment group when regional markets with insiders are excluded from the control group.

Second, we further adjust the definition of the control group to identify upward-pricing pressure induced by the merger. More precisely, we adjust the control group such that it contains all local markets where at least one of the insiders operates (see Figure 1(c)). This is the opposite of the aforementioned case in which control groups are defined as local markets with outsiders only. The specification identifies price increase due to an increase in market power. Since insiders' stores are present in both treatment and control group, global efficiency gains are canceled out. Identification of this effect hinges on the assumptions (i) that retailers employ a national bargaining strategy on the supply side instead of leaving negotiations to local merchants and (ii) there is equal pass-through of efficiency gains in treatment and control group. While we feel confident to assume the former given anecdotal evidence discussed above, the latter assumption might be violated. However, if there are indeed local efficiency gains in markets with both acquirer's and target's stores, this would imply relatively lower prices in the treatment group, and our estimates of price increases due to increased market power can be interpreted as a lower bound.

Finally, we investigate the pass-through of efficiency gains. In this specification, our control group consists of markets with outsiders only, while the treatment group contains markets in which one but not both of the merging parties were active before the merger (Figure 1(d)). Since, in this specification, we compare markets without pre-merger overlap of acquirer and target to a control group of markets without insiders, there are no systematic changes in local competition that affect the estimates. Hence, relative price changes stem from changes in outsiders' responses to price decreases due to insiders' efficiency gains.

Figure 2.3 summarizes the different definitions of treatment and control markets for our various settings. Markets of type A are those where we expect efficiency gains and market power effects to be present. Markets of type B are markets that presumably are affected by efficiency gains but not by changes in market power since there is no pre-merger overlap of acquirer and target in these markets. Markets of type C contain outsiders only and thus we expect neither market power nor efficiency effects to play an important role. Concentration effects are identified by comparing type A markets to type B markets. Efficiency gains are identified by a comparison of type B markets with type C markets. Figures 3–6 show pre-merger price trends for the different market types. For all settings, pre-merger price trends are similar between treatment and control group for all product categories.

It is worth noting that the year of the financial crisis in 2008 is covered in our estimation sample. However, we believe that this incidence does not drive our results for three reasons. First, its impact on Germany's economy lower relative to other European countries.[12] Second, although it might be that the financial crisis impacted treatment and control regions to a different extent, figure 2.7 shows that economic indicators evolved similarly within both types of regions. Third and finally, despite the fact that economic indicators do not impact treatment and control group in a different way, we nonetheless include unemployment rate and household income as control variables into our regression analysis to account for the fact that they might have a potential impact on the results conditional on other control variables and fixed effects.

2.4.2 Empirical implementation

For our baseline specification, we use a simple difference-in-differences (DiD) estimator to analyze the effects of the merger on regional consumer prices. The assignment of local markets to treatment and control group in this baseline specification corresponds to the definition in Figure 2(a). In section 5.2, we present further results on varying local market types (Figure 2(b)–Figure 2(d). The baseline specification

[12] Both national and international newspapers reported that Germany's strategies to recover from the financial crisis worked quite well. See, for instance, `http://www.nytimes.com/2010/08/14/world/europe/14germany.html` or `http://www.spiegel.de/international/business/global-debt-disaster-what-the-financial-crisis-means-for-germany-a-779306.html`, both accessed November 12, 2017.

estimates:

$$ln(p_{igjt}) = \alpha_{igj} + \theta\, post_t \times MA_g + \delta_t + [x'_{gt}\beta + \eta_{it} + \omega_{kt}] + \varepsilon_{igjt} \quad (2.1)$$

where $ln(p_{igjt})$ denotes the logarithmic product price set by retail chain i in regional market g (defined at the county level, e.g. Cologne), for brand j at quarterly time period t. Since both merging retailers have different product portfolios pre- and post-merger, we calculate mean prices for both retailers' product portfolios in the estimation sample. However, since there might be composition effect, we implement a robustness check, where we treat branded products of the merging retail chains like a product of one retailer in markets with pre-merger overlap of $R1$ and $R2$. Private labels sold by the target retail chain in re-branded stores will not enter the estimation since they disappear in the post-merger period.

α_{igj} is a retailer-market-product fixed effect, $post_t$ takes on a value of one in all post-merger periods, MA_g is a dummy variable indicating regions affected by the merger and δ_t denotes a full set of time dummies. In some specifications, we add the terms in brackets: x'_{gt} controls for time-varying demand heterogeneity at the regional level through changes in average income, population density and unemployment. η_{it} denotes retail chain-time fixed effects which control for national-wide price changes across retailers. Note that these also capture any concentration and efficiency effects of the merger which do not vary across regions. ω_{kt} controls for overall price changes in product categories (k) across time.[13] Finally, ε_{igjt} is an error term. Our main coefficient of interest is the DiD-parameter θ which indicates differences between adjustments of consumer prices within retailer-products across regions at the time of the merger. Since our dependent variable is retailer-product-region specific, while our treatment indicator only varies across regions within years, we compute standard errors that are clustered at the region level.

We extend our baseline specification in several dimensions. First, we analyze heterogeneous effects across private label products and national brands by estimating the following specification:

$$ln(p_{igjt}) = \alpha_{igj} + \theta_1 post_t \times MA_g + \theta_2 post_t \times MA_g \times PL_j + \delta_t + [x'_{gt}\gamma + \eta_{it} + \omega_{kt}] + \varepsilon_{igjt} \quad (2.2)$$

[13]To capture product-specific trends, we had to aggregate products j to product categories k since we do not observe purchases of all products in all regions.

where PL_j takes on a value of one if product j is sold under a private label. In this specification, θ_1 measures the treatment effect for products sold under national brand names, while θ_2 indicates differences between price effects for private label products relative to those for national brands.

Next, we extend equation (2.1) to allow for heterogeneous retail chain-specific treatment effects:

$$ln(p_{igjt}) = \alpha_{igj} + \varphi_1 post_t \times MA_g + \varphi_2 post_t \times MA_g \times DC_i + \delta_t + [x'_{gt}\pi + \eta_{it} + \omega_{kt}] + \varepsilon_{igjt} \tag{2.3}$$

where DC_i takes on a value of one for discounters and φ_2 measures differences in price changes between supermarkets and discounters. We also use a similar specification in which we replace DC by a dummy variable for insiders to distinguish between effects on merging parties and non-merging competitors.

In another extension of our baseline model, we investigate whether price increases are more likely to occur in markets with high expected changes in concentration. For this purpose, we follow Dafny et al. (2012) and construct the predicted change in the Herfindahl-Index (HHI) induced by the merger:

$$\Delta simHHI_{gk} = 2 \times Acqshare_{gk} \times Tarshare_{gk} \tag{2.4}$$

where $Acqshare_{gk}$ and $Tarshare_{gk}$ denote the pre-merger market shares for region g and product category k of acquirer and target, respectively. For instance, if acquirer and target have pre-merger market shares of 10% each, HHI, the sum of squared market shares, would be expected to change by 0.02. We use this predicted change in concentration to test the hypothesis that price increases are more likely to occur in regional markets with substantial changes in retail concentration in the following equation:

$$ln(p_{igjt}) = \alpha_{igj} + \tau_1 post_t \times MA_g + \tau_2 post_t \times MA_g \times \Delta simHHI_{gk} + \delta_t + [x'_{gt}\kappa + \eta_{it} + \omega_{kt}] + \varepsilon_{igjt} \tag{2.5}$$

In this specification, τ_1 estimates the effect of the merger on prices that is independent of initial market shares of acquirer and target. τ_2 captures heterogeneity of treatment effects with respect to variation in expected regional retail market concentration.

Finally, we analyse whether remedies imposed by the German cartel office, which required the sale of several retail stores to a competitor in regional markets with high pre-merger market shares of acquirer and target, had the desired effects. For this purpose, we estimate:

$$ln(p_{igjt}) = \alpha_{igj} + \lambda_1 post_t \times MA_g + \lambda_2 post_t \times MA_g \times Rem_g + \delta_t + [x'_{gt}\nu + \eta_{it} + \omega_{kt}] + \varepsilon_{igjt} \tag{2.6}$$

where Rem_g takes on a value of one if remedies were imposed in market g and MA_g indicates regions affected by the merger which includes regions with and without remedies. Thus, λ_1 measures the estimated impact of the merger in non-remedy regions and λ_2 indicates whether there are different effects in regional markets in which remedies have been imposed.

2.5 Results

This section presents the empirical findings. We first discuss the baseline results in section 5.1 where we analyze average and heterogeneous treatment effects by comparing relative price changes of the treatment group with a control group focusing on local effects and abstracting from national pricing strategies and efficiency gains. We introduce and discuss heterogeneous treatment effects for varying market power, remedy and non-remedy regions, and discounters and supermarkets. The next section, 5.2, discusses the results of a setup in which market concentration effects and efficiency gains are identified using different definitions of treatment and control groups along with the net effect accounting for national strategies. Finally, we present a series of robustness tests in section 5.3.

2.5.1 Baseline results

Table 4.4.2 summarizes the baseline results which are based on the assumption that there are no national-wide price effects of the merger. Regions affected by the merger might be systematically different from regions without stores operated by the merging parties. To account for this possible confounding, we include fixed effects for region-retailer-brand, retailer-time, and category-time. We also control for time-varying variables at the regional level including population density, mean income, mean age, unemployment rate and the average number of children per household. Thus, we assume that assignment to treatment and control group is random conditional on our rich set of covariates.

Models 1 shows a positive significant treatment effect for the baseline specification. Model 2 includes an interaction term with the expected change in concentration, ΔHHI. Markets with relatively high pre-merger market shares of the merging parties should be affected relatively more by the merger if prices respond to changes in market concentration. The positive significant interaction term confirms this intuition. The interaction effect is also economically significant. For instance, for a market in which $R1$ and $R2$ hold pre-merger market shares of 30% and 15%, respectively, the model predicts prices to increase by 1%. For the highest value of ΔHHI in the sample, which corresponds to a region in which acquirer's and target's pre-merger market shares were close to 50% each, the estimates indicate that prices increased by about 4.24% due to the merger.

Model 3 extends model 1 through an interaction term between treatment regions and discounters. The average treatment effect for non-discounters is 0.77% and highly statistically significant, indicating that supermarkets have, on average, slightly raised prices due to the merger. The insignificant result for discounters indicates that this retail format is perceived as a limited substitute for supermarkets. Model 4 combines Model 2 and 3 and includes separate interaction terms with expected changes in market concentration for discounters and supermarkets. The results indicate that supermarkets, but not discounters, increased prices in regions with high expected change in retail concentration. For instance, the model predict a price increase of 1.9% in a market where acquirer and target had pre-merger market shares of 30% and 15% and a maximum price increase of 7.6% in a hypothetical market where $R1$ and $R2$ both have 50% market share before the merger. Models 5 and 6 are based on the same estimation equations as models 1 and 3, respectively, but are based on a treatment group that consists of remedy regions only and excludes non-remedy regions that contain stores of both $R1$ and $R2$. The estimated effect in regions where remedies have been imposed is larger than the average effect in the baseline specification. As discussed in section 2.2.1, remedy regions consist of relatively highly concentrated local markets (Bundeskartellamt, 2007). Our results indicate that remedies imposed were not sufficient to offset anti-competitive effects of the merger.

2.5.2 Market concentration, efficiency gains, and net effect with national strategies

In this subsection, we provide further evidence for heterogeneous treatment effects which are related to local market structure. We first present the net effect of market concentration and efficiency gains, which allow for national-wide effects of the merger. As outlined in section 4.1.2, identification of net effect and efficiency gains is based on the assumption that outsiders adjust their prices to insiders' cost savings, for which we provide evidence in table B2.

Model 7 in Table 2.5 identifies the net effect of market concentration and efficiency gains taking into account national strategies. In this specification, the treatment group consists of outsiders in markets with pre-merger overlap of acquirer and target, while the control group includes outsiders in markets without any presence of the merging parties. Differences between treatment and control group are therefore likely to stem from both changes in market concentration and from outsiders' reaction to efficiency gains and corresponding price adjustments of insiders. The estimated average effect is negative but non-significant. Model 8 includes an interaction term with the expected change in market concentration which again indicates that price increases are more likely in regions with higher pre-merger market shares of acquirer and target.

In specification 9, we restrict the control group to regions in which exactly one of the merging retailers was active. If we expect efficiency gains to play a role, the estimated coefficient should be lower in magnitude in the net effect specification (model 7) compared to the market power specification (model 9) where we identify price effects that stem from an increase in market concentration net of national-wide cost savings. Intuitively, insiders in both treatment and control group can benefit from lower costs due to national-wide efficiency gains or higher bargaining power after the merger. If these cost savings are passed-on to consumers, outsiders in both treatment and control group are likely to adjust their prices as well. At the same, market concentration changes systematically in the treatment group but not in the control group. As expected, there is a larger treatment effect in this specification than in the net effect model. Prices in regions affected by market concentration increase, on average, by 0.55%. Models 10 includes an interaction term with the expected change in concentration which is again significant and positive. Hence, we indeed find support for the hypothesis of price increases in regions where changes in market concentration are likely to dominate efficiency gains from the merger. This result is in line with several studies on retailing markets which indicate that higher seller concentration (less competition) is associated with higher prices (von Ungern-Sternberg, 1996; Dobson and Waterson, 1997). Our results are also in line with Dafny et al. (2012) and Ashenfelter et al. (2015) who find that prices increased substantially in markets with high predicted changes in market concentration.

In model 11 and 12, we use the same control group as in model 9 and 10, but the treatment group now consists of markets in which only one of the merging retailers was active prior to the merger. In this specification, retailers in the treatment group operate in a market where the parties potentially experience a reduction in costs which might induce price reactions from both insiders and outsiders, but there is no systematic change in market concentration in either treatment or control group. This allows us to identify downward pressure on prices due to national-wide efficiency gains. The treatment effect is (weakly significantly) negative, indicating a decrease in relative prices of 1.0% on average. Intuitively, we would expect that the pass through of lower costs to consumer prices is higher in competitive markets. Model 12 includes an interaction term with the pre-merger *level* of the HHI.[14] The base effect increases in absolute terms and becomes more significant. The interaction term is positive and statistically significant which indicates lower price reductions induced by cost savings in ex ante more concentrated markets. Hence, we find support for pass-through of cost savings in markets where merging firms were likely to benefit from cost savings, a result which is consistent with the argument of Williamson (1968). It seems noteworthy that this specification does not allow identifying the magnitude of efficiency gains since we measure the pass-through of cost-savings on consumer prices. Moreover, although we quantify the decrease in prices for consumers of certain markets, we are not able to identify the source of these price changes, which presumably stem from efficiency gains or increased bargaining power.

Price reactions to the merger presumably take time as retailers have to adjust to the new industry structure. This is particularly true in the case of cost savings which might not be realized immediately after the merger. Table 2.6 shows heterogeneous treatment effects across time for the market concentration, the net, the efficiency, and the baseline specification. The market power specification shows that prices already increased in the year 2009 but to a larger extent in 2010, statistically and

[14] Note that the predicted change in market concentration, ΔHHI, is zero for all markets in this specification.

economically. The efficiency specification shows that the average effect is driven by the year 2010 indicating that efficiency gains took one year to be realized.

Table B3 in the appendix shows interaction terms between discounters and indicators for treatment regions in columns 1 and 2. Time heterogeneous treatment effects for discounters and non-discounters are depicted in columns 3–5. The first column indicates that the market power effect is mainly driven by supermarkets, whereas results are less clear for responses to potential cost savings. Columns 3–5 indicate that treatment effects are more pronounced in 2010 compared to 2009.

Our study contributes to understanding the ambiguous results produced by an increasing number of studies on ex-post merger analysis in the retail sector. For instance, Allain et al. (2017) report a significant price increase after a merger among French retailers while Argentesi et al. (2016) find no significant price changes within product categories after mergers in the Dutch retail market and Hosken et al. (forthcoming) analyze 14 retail mergers and find that prices decreased after some mergers while they increased or remained unchanged in other cases. Our results indicate that the mixed results might be explained by a combination of efficiency gains and market concentration effects which are not uniform across local markets and, to some extent, by heterogeneous responses of supermarkets and discounters.

2.5.3 Robustness checks

In this section, we discuss the results of various robustness checks. First, we use alternative measures of aggregate prices. For instance, we replace mean prices with median prices and construct prices weighted by the number of purchases. Further, we treat the merging parties as a single firm in both pre and post merger periods. Second, we remove neighboring markets from the control group to account for possible spillovers. Third, we employ a propensity score re-weighting estimator. Finally, we perform placebo tests in order to check whether our results can be explained by heterogeneity in unobserved trends.

In the baseline specification, we use average unit prices across purchases within a region. Estimates using mean prices could be potentially affected by outliers. To mitigate this concern, Table B4 shows results using median prices instead of mean prices. In the baseline specification, we do not account for quantities purchased at a certain price when we construct mean prices at the regional level. Table B5 shows results for weighted average prices where we calculate prices as the ratio between sales and quantity purchased. Our main results, obtained in models 1–4 of Table 4.4.2 are robust towards these alternative measures of local prices.

The estimates could be driven by composition effects, for instance, if the merging retailers—instead of raising the prices within products—adapt their portfolio towards more expensive products. To assess whether such a strategy affects our results, we treat the merging parties as a single retailer and re-implement our DiD analysis. Results of this specification are shown in Table B6. To further investigate variety effects, we test in Table B7 if retailers adapt the number of products sold post-merger. The results show that there is a change in variety of inside discounters. The effect is more pronounced in markets with a larger expected change in market concentration. However, the price changes cannot be explained by this change in variety. The positive price effect is driven by insider and outsider supermarkets, whereas variety changes are driven by inside discounters. The price changes due to efficiency gains are identified by outsiders but the variety results indicate that outsiders do not change the variety.

Another potential concern relates to the definition of local markets and spillovers to regions which are part of the control group. If retailers in neighboring markets, assigned to our control group, adjust their prices in response to prices in treatment regions, they do not provide an accurate estimate of the counterfactual. To address this problem, we remove all markets from the control group, which are within a 15 km radius of markets in the treatment groups, in a further robustness check. Table B8 shows specifications which correspond to models 1–4 of Table 4.4.2. In this specification, the average treatment effect increases compared to the baseline specification. Heterogeneous effects indicated by interaction terms are of similar magnitude as in the baseline specification. In another specification we remove all cities from the sample. Table B9 shows the results of this robustness check. The average effect disappears, but the effect in the highly concentrated markets remain.

As a further robustness check, we use a propensity score re-weighting estimator. This estimator assigns higher weight to control markets that are more similar to markets in the treatment group. Matching has been performed on average pre-merger growth rates of prices and demographics. Table B10 shows the results of the matching procedure indicating that the variables are balanced between treatment and control group after reweighting. Table B11 shows regression results in which observations in the treatment group are assigned a weight equal to 1 and control markets are assigned a weight of p/(1-p), where p is the estimated probability of being treated. The main results hold in this set up suggesting that the treatment effect is not driven by fundamental differences in characteristics of treatment and control markets.

In addition to these alternative specifications, we performed two placebo tests. In the first test, we randomly assign a market to the treatment group but use the actual time of the merger to define pre- and post treatment observations. We perform this procedure ten times and estimate model 3 of Table 4.4.2 for each random draw. Table B12 presents the results of this test. None of the draws shows significant results indicating that our results with the correct definition of the treatment group have not been obtained by chance.

A potential concern in DiD regressions is that estimated effects picks up heterogeneous trends in treatment and control group. To check whether heterogeneous trends can explain our results, we vary the treatment time definition by assigning it to dates before the actual merger took place and discard all observations in the merger year 2007 and post-merger observations. Thus, we use the correct regional definition of treatment and control groups but restrict the sample to pre-merger years. If there are systematically different pre-merger time trends between treatment and control group, this should result in a significant estimated treatment effect in this placebo test. Table B13 shows the result of three different definitions of the treatment time period. All specifications yield insignificant estimated treatment effects indicating that pre-merger time trends are unlikely to affect our results. Note, we can not conduct the placebo tests adding the interaction of treatment effect with the expected change in the HHI since that variable is defined for regions in the treatment group solely.

2.6 Conclusion

In this paper, we analyze the impact of a retail merger in which a supermarket acquired a soft discounter on consumer prices in Germany. We exploit geographical and time series variation in price setting by retail chains, and the fact that both acquirer and target were not active in all regional markets before the merger, to estimate the causal effect of the retail merger on prices. We find a small average prices increase and considerable heterogeneity of merger effects. The estimated effects increase significantly with predicted changes in market concentration and amount to about 4% in markets with the highest pre-merger market shares of acquirer and target. Our results indicate that the treatment effect is driven by supermarkets rather than discounters and that average price reactions are more pronounced in remedy regions where the acquirer had to sell target's stores to an outsider. We also provide evidence that efficiency gains have partly offset price increases due to market power in some regions. It seems that the merger even resulted in lower prices in regions that were potentially affected by national-wide efficiency gains but not by a change in local market concentration. We find the largest relative price increases when we compare regions with ex ante predicted changes in markets concentration to regions that are not affected by a change in market structure but are presumably similarly affected by national-wide cost savings of the merged entity. This cross-regional heterogeneity of merger effects can potentially explain the ambiguous results of previous ex post merger evaluations in the retail sector.

Our results have important implication for economic policy and competition authorities in particular. First, they indicate that increasing retail concentration in general is a potential concern for consumer welfare as it is likely to lead to higher prices. Second, our results suggest that price increases after mergers can be predicted by pre-merger market shares of acquirer and target. Moreover, our results indicate that cost savings induced by retail mergers can be significant and partly offset anti-competitive effects. Finally, we also provide evidence that remedies imposed by competition authorities were not sufficient to offset anti-competitive effects.

There are several possible directions for future research. First, we would like to develop a structural model that is able to disentangle the channels of price changes in more detail and to decompose cost savings into efficiency gains and changes in bargaining power. Second, it would be valuable to analyze whether cross-regional variation in price changes can be predicted by merger simulation tools adjusted to the retail sector. Finally, given the increased availability of consumer and store level data, it will be interesting to see to which extent our results apply to different merger cases in various countries.

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2.7 Figures

Figure 2.1: Analysis of Variance in Local and National Component

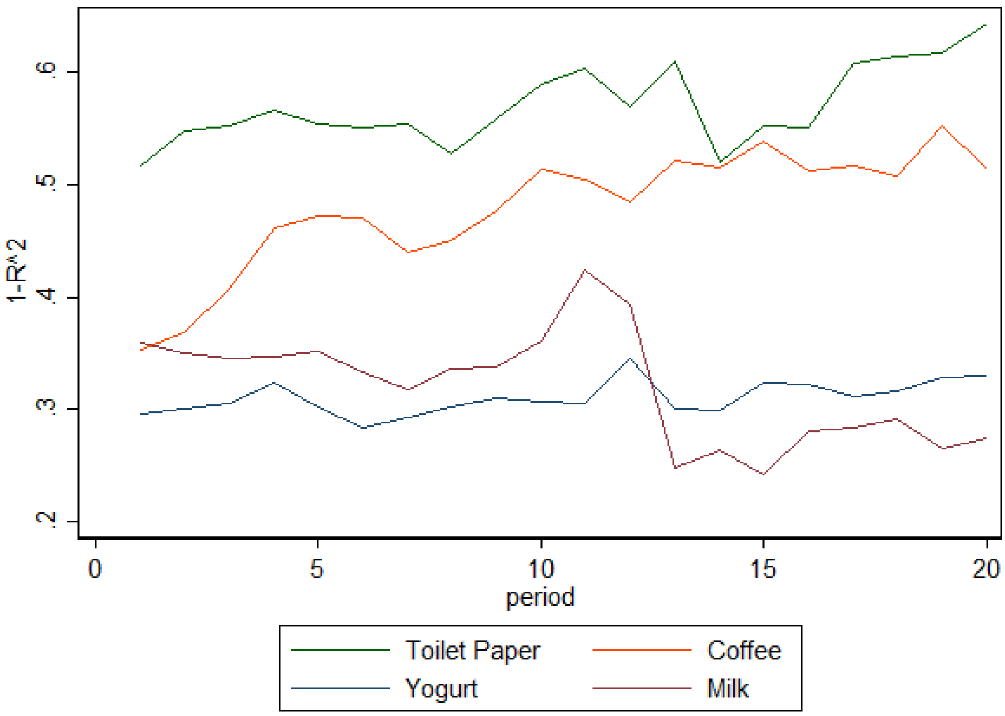
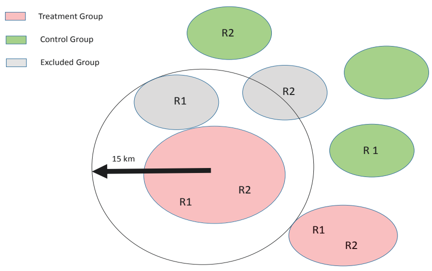
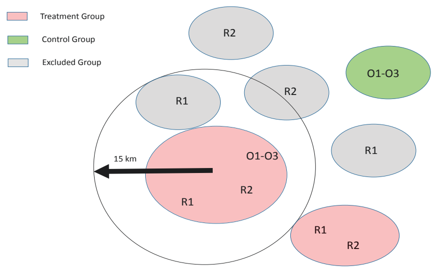


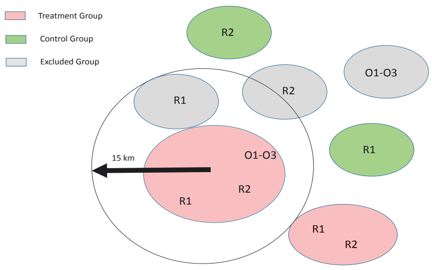
Figure 2.2: Definition of Treatment and Control Groups



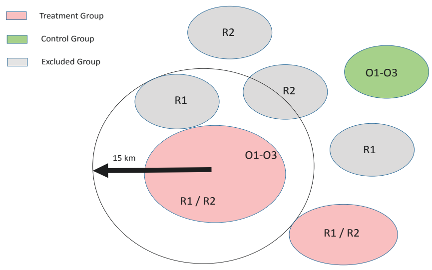
(a) Specification 1: Baseline



(b) Specification 2: Net Price Effect



(c) Specification 3: Concentration Effects



(d) Specification 4: Efficiency Gains Pressure

Figure 2.3: Parallel Trends over Categories in Baseline Specification

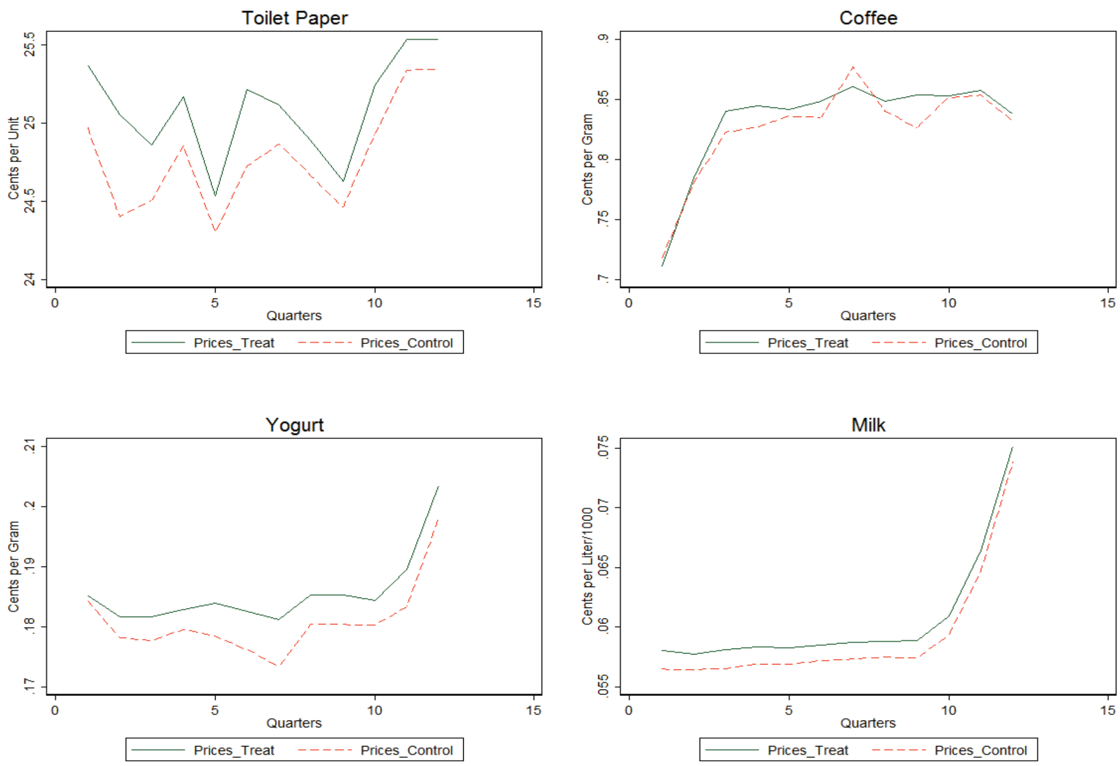


Figure 2.4: Parallel Trends over Categories in Net Effect with National Strategies Specification

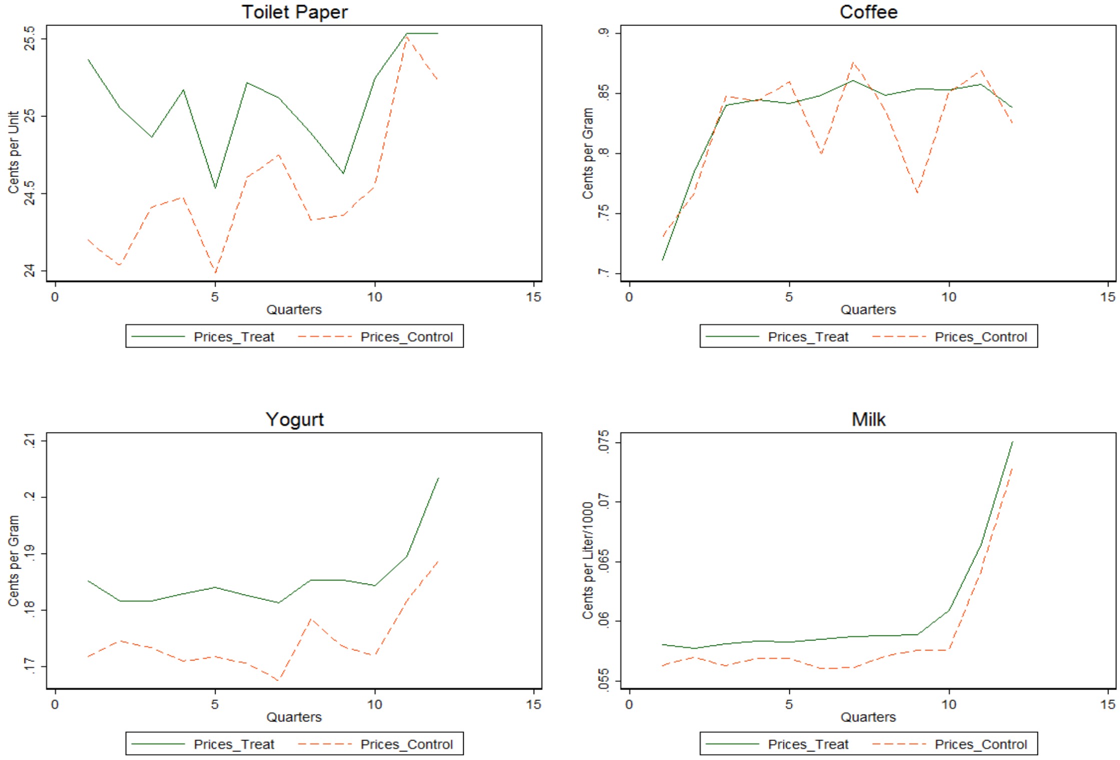


Figure 2.5: Parallel Trends over Categories in Market Concentration Specification

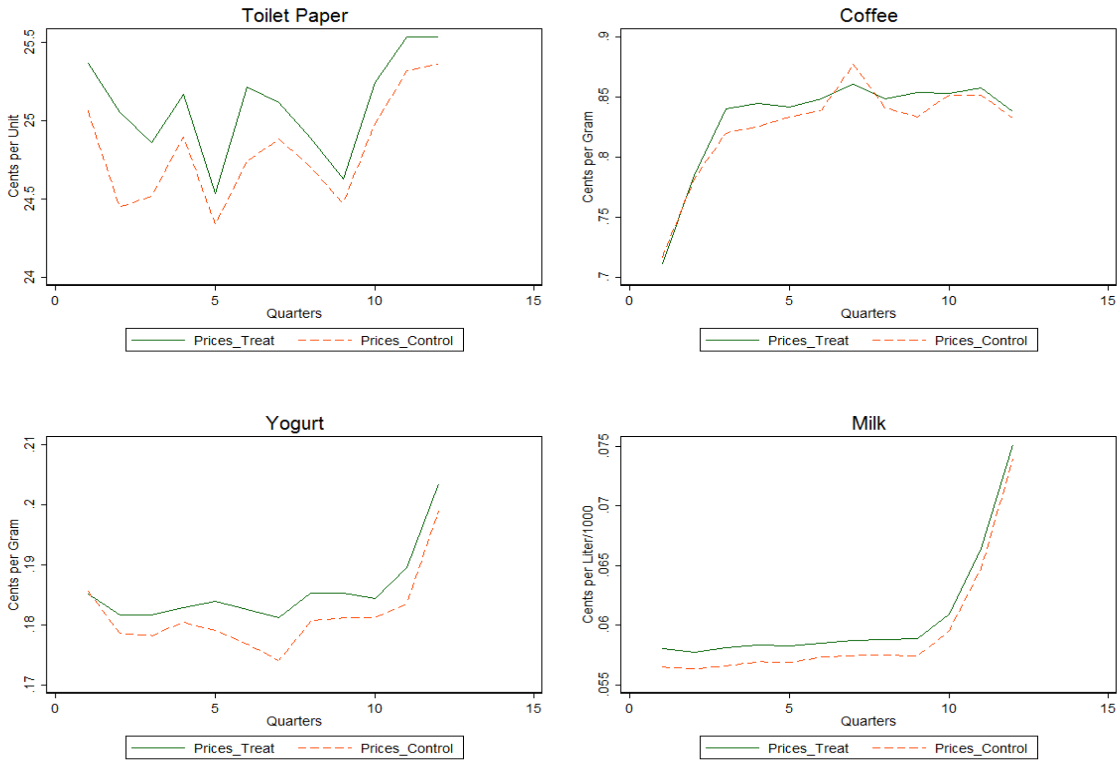


Figure 2.6: Parallel Trends over Categories in Efficiency Gains Specification

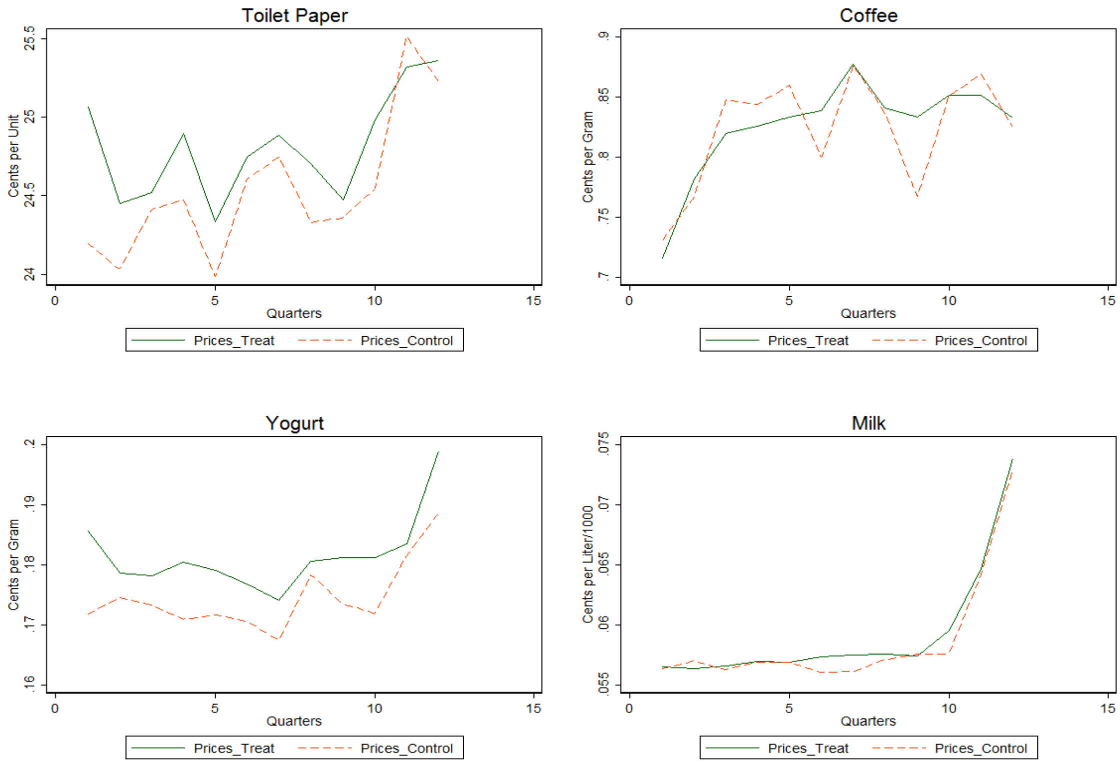
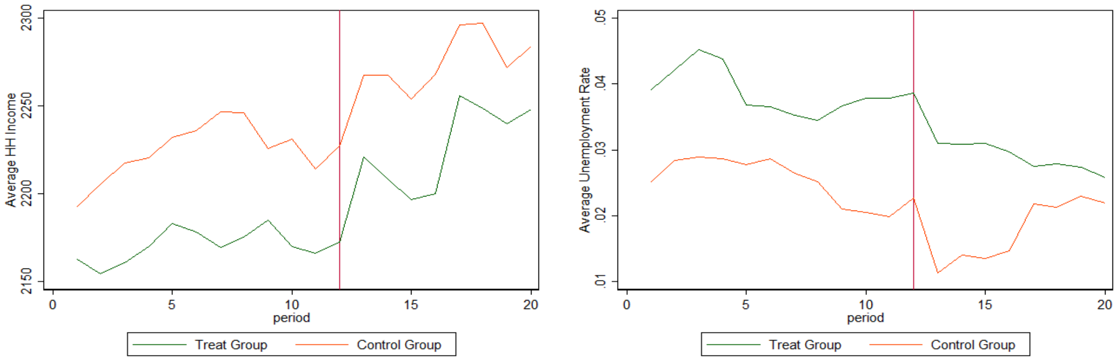


Figure 2.7: Average Income and Unemployment Rate



2.8 Tables

Table 2.1: Market Shares in Percent

		Supermarkets		Discounters	
		Pre	Post	Pre	Post
Toilet Paper	National Brands	15.48	12.76	2.36	2.31
	Private Labels	26.67	29.76	55.49	55.17
Coffee	National Brands	52.92	49.82	16.49	22.28
	Private Labels	3.72	2.57	26.87	25.33
Yogurt	National Brands	37.03	33.63	9.26	11.02
	Private Labels	6.61	10.47	47.11	44.89
Milk	National Brands	15.97	12.45	2.55	2.17
	Private Labels	27.63	30.53	53.86	54.84

Notes: Market shares were calculated using revenue data from all shopping trips conducted by the representative panel members selected by the Gesellschaft für Konsumforschung (GfK). Shares were calculated from sales data from January 2005 through December 2010 without the year 2008. a market here is defined by pre and post merger and product category.

Table 2.2: Average Prices by Product Category

		Supermarkets		Discounters	
		Pre	Post	Pre	Post
Toilet paper	Control	25.82	28.31	23.64	26.65
	Treat	25.92	28.53	23.48	26.74
Coffee	Control	92.84	101.6	70.82	74.05
	Treat	92.15	100.6	72.45	75.27
Yogurt	Control	20.3	22.33	15.03	16.81
	Treat	20.23	22.27	14.99	16.91
Milk	Control	6.31	6.32	5.47	5.3
	Treat	6.2	6.35	5.47	5.27

Notes: The table reports average prices by retail format, product category, and treatment status. Average prices are calculated from the observed purchases by consumers in our sample. According to the respective order, pricing units are: cents per unit, cents per 100 gram, cents per 100 gram, cents per unit, cents per 100 ml.

Table 2.3: Market Types

Type	Players	Concentration	Efficiency	Frequency
A	R1 **and** R2, O	X	X	1816
B	R1 **or** R2, O	-	X	1417
C	O	-	-	424

Table 2.4: Baseline Results

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Treat	0.0047**	0.0039*	0.0077***	0.0063**	0.0042	0.013***
	(0.0021)	(0.0021)	(0.0030)	(0.0030)	(0.0035)	(0.0048)
Treat × Δ*HHI*		0.077**		0.14***		
		(0.039)		(0.048)		
Treat × *DC*			-0.0061*	-0.0046		-0.016***
			(0.0035)	(0.0036)		(0.0056)
Treat × *DC* × Δ*HHI*				-0.15**		
				(0.063)		
Region	all	all	all	all	remedy	remedy
FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls						
N	816103	816103	816103	816103	289771	289771
R^2	0.994	0.994	0.994	0.994	0.994	0.994

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. DC is a dummy variable taking value 1 if the supermarket is a discounter. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table 2.5: Disentangling Price Effects

Effect	Model 7 Net	Model 8 Net	Model 9 M Power	Model 10 M Power	Model 11 Efficiency	Model 12 Efficiency
Treat	-0.0041	-0.0048	0.0055**	0.0048**	-0.0100*	-0.016***
	(0.0054)	(0.0054)	(0.0022)	(0.0022)	(0.0056)	(0.0058)
$Treat \times \Delta HHI$		0.073*		0.076*		
		(0.039)		(0.039)		
$Treat \times HHI$						0.022***
						(0.0059)
FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls						
N	627463	627463	794766	794766	209947	209947
R^2	0.994	0.994	0.994	0.994	0.996	0.996

Notes: The dependent variable are log mean prices at the region-retailer-brand level. In columns 1, 2, 3, 4 Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. In columns 5 and 6 Treat takes value 1 post merger in markets where only one of the merging parties is present. In columns 3 and 4 the control group consists of markets with exactly one of the merging parties present. In all other columns the control group are markets where no merging party is active. ΔHHI is the expected change in market concentration. HHI is the HHI index in levels. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table 2.6: Disentangling Time Heterogeneous Price Effects

Effect	Model 7 M Power	Model 9 Net	Model 11 Efficiency	Model 1 Baseline
Treat × *2009*	0.0041*	-0.0017	-0.0063	0.0036*
	(0.0022)	(0.0055)	(0.0057)	(0.0022)
Treat × *2010*	0.0071***	-0.0067	-0.014**	0.0058**
	(0.0025)	(0.0061)	(0.0064)	(0.0024)
FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls				
N	794766	627463	209947	816103
R^2	0.994	0.994	0.996	0.994

Notes: The dependent variable are log mean prices at the region-retailer-brand level. In columns 1, 2 and 4 Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. In column 3 Treat takes value 1 post merger in markets where only one of the merging parties is present. In column 1 the control group consists of markets with exactly one of the merging parties present. In columns 2 and 3 the control group are markets where no merging party is active. The control group in the baseline scenario are all markets without overlap of the merging parties. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

2.9 Appendix

Table B1: Regional Characteristics

Population Density	0.12***	0.11***	0.010***	0.0092***	0.0096***
	(0.0098)	(0.0096)	(0.0014)	(0.0012)	(0.0012)
Unemployment Rate	0.066	0.066*	-0.0090	-0.0099	-0.0098
	(0.042)	(0.039)	(0.0078)	(0.0077)	(0.0076)
Average Children	-0.0046	-0.0069	-0.0070***	-0.0072***	-0.0072***
	(0.0096)	(0.0092)	(0.0018)	(0.0018)	(0.0017)
HH Income	0.0015	0.0016	0.0054***	0.0067***	0.0066***
	(0.0090)	(0.0090)	(0.0017)	(0.0016)	(0.0016)
Age	-0.47	0.00072	-0.0011***	-0.0010***	-0.99***
	(0.73)	(0.71)	(0.14)	(0.13)	(0.13)
HHI	0.48***	0.58***	-0.010***	-0.0084**	-0.0092***
	(0.034)	(0.031)	(0.0040)	(0.0035)	(0.0035)
Time FE	yes	yes	yes	no	no
Retailer FE	no	yes	yes	yes	no
Brand FE	no	no	yes	yes	yes
Category-Time FE	no	no	no	yes	yes
Retailer-Time-PL	no	no	no	no	yes
N	816103	816103	816103	816103	816103
R^2	0.005	0.092	0.979	0.986	0.986

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B2: Heterogeneous Effects for Insiders and Outsiders

	Model 1	Model 2	Model 3	Model 6
Treat × *Outsider*	0.0049**	0.0045**	0.0082***	0.011**
	(0.0021)	(0.0021)	(0.0031)	(0.0051)
Treat × *Insider*	0.0032	0.0013	0.0065	0.022*
	(0.0035)	(0.0036)	(0.0045)	(0.012)
Treat × *Outsider* × ΔHHI		0.055		
		(0.061)		
Treat × *Insider* × ΔHHI		0.10**		
		(0.045)		
Treat × *Outsider* × *DC*			-0.0062*	-0.016***
			(0.0036)	(0.0060)
Treat × *Insider* × *DC*			-0.0082	-0.019
			(0.0064)	(0.017)
Region	all	all	all	remedy
FEs for Region-Retailer-Brand, Retailer-Time, Category-Time + Regional Controls				
N	816103	816103	816103	289771
R^2	0.994	0.994	0.994	0.994

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. Insider and Outsider are dummy variables taking the value 1 post merger for inside and outside firms, respectively. DC is a dummy variable taking the value 1 for products sold at a discounter. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B3: Disentangling Time Heterogeneous Price Effects for DC and Non-DC

Effect	Model 13 MP	Model 14 Efficiency	Model 13 MP	Model 14 Efficiency	Model 15 Baseline
Treat × *NON-DC*	0.0085***	-0.010			
	(0.0030)	(0.0093)			
Treat × *DC*	0.0026	-0.0098*			
	(0.0026)	(0.0058)			
Treat × *NON-DC* × *2009*			0.0064*	-0.0077	0.0059*
			(0.0033)	(0.0092)	(0.0032)
Treat × *DC* × *2009*			0.0018	-0.0051	0.0014
			(0.0027)	(0.0065)	(0.0026)
Treat × *NON-DC* × *2010*			0.011***	-0.013	0.0098***
			(0.0035)	(0.011)	(0.0034)
Treat × *DC* × *2010*			0.0035	-0.015**	0.0020
			(0.0029)	(0.0061)	(0.0028)
N	794766	209947	794766	209947	816103
R^2	0.994	0.996	0.994	0.996	0.994

Notes: The dependent variable are log mean prices at the region-retailer-brand level. In columns 1, 3 and 5 Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. In column 2 and 4 Treat takes value 1 post merger in markets where only one of the merging parties is present. In column 1 and 3 the control group consists of markets with exactly one of the merging parties present. In columns 2 and 4 the control group are markets where no merging party is active. The control group in the baseline scenario are all markets without overlap of the merging parties. DC is a dummy variable taking the value 1 for discounters. NON-DC is a dummy variable taking value 1 for non discounters. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B4: Baseline Results: Median Prices

	Model 1	Model 2	Model 3	Model 4
Treat	0.0041*	0.0035	0.0068**	0.0054*
	(0.0022)	(0.0022)	(0.0030)	(0.0031)
Treat $\times$ ΔHHI		0.067*		0.13***
		(0.039)		(0.049)
Treat $\times$ *DC*			-0.0053	-0.0038
			(0.0037)	(0.0038)
Treat $\times$ *DC* $\times$ ΔHHI				-0.15**
				(0.064)
	FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls			
N	816103	816103	816103	816103
R^2	0.993	0.993	0.993	0.993

Notes: The dependent variable are log median prices at the region-retailer-brand level. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B5: Baseline Results: Weighted Prices

	Model 1	Model 2	Model 3	Model 4
Treat	0.0045**	0.0039*	0.0077***	0.0064**
	(0.0021)	(0.0021)	(0.0030)	(0.0030)
Treat $\times$ Δ*HHI*		0.068*		0.13***
		(0.039)		(0.048)
Treat $\times$ *DC*			-0.0063*	-0.0049
			(0.0035)	(0.0036)
Treat $\times$ *DC* $\times$ Δ*HHI*				-0.14**
				(0.063)
	FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls			
N	816103	816103	816103	816103
R^2	0.994	0.994	0.994	0.994

Notes: The dependent variable are log mean prices weighted by the number of purchases at the region-retailer-brand level. Neighboring markets are dropped from the sample. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B6: Baseline Results: Mean Prices for Merging Discounters

	Model 1	Model 2	Model 3	Model 4
Treat	0.0026	0.0023	0.0076**	0.0062**
	(0.0021)	(0.0021)	(0.0030)	(0.0030)
Treat × Δ*HHI*		0.031		0.14***
		(0.039)		(0.048)
Treat × *DC*			-0.0098***	-0.0076**
			(0.0035)	(0.0036)
Treat × *DC* × Δ*HHI*				-0.22***
				(0.067)
	FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls			
N	814780	814780	814780	814780
R^2	0.994	0.994	0.994	0.994

Notes: The dependent variable are log mean prices at the region-retailer-brand level. The discounter of the merging parties are treated as one retailer. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B7: Variety Results

Treat	-0.0018	0.0013	-0.00042	0.0011	
	(0.0019)	(0.0029)	(0.0020)	(0.0029)	
Treat × *DC*		-0.0060		-0.0030	
		(0.0039)		(0.0039)	
Treat × ΔHHI			-0.14**	0.024	
			(0.062)	(0.081)	
Treat × *DC* × ΔHHI				-0.32***	
				(0.100)	
Treat × *Insider*					-0.0036
					(0.0049)
Treat × *Outsider*					0.0029
					(0.0032)
Treat × *Insider* × *DC*					-0.060***
					(0.0094)
Treat × *Outsider* × *DC*					0.0033
					(0.0043)
Treat × *Insider* × ΔHHI					0.11
					(0.093)
Treat × *Outsider* × ΔHHI					-0.091
					(0.11)
Treat × *Insider* × *DC* × ΔHHI					-0.51***
					(0.12)
Treat × *Outsider* × *DC* × ΔHHI					0.18
					(0.12)
FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls					
N	814794	814794	814794	814794	814794
R^2	0.500	0.500	0.500	0.500	0.501

Notes: The dependent variable are log mean number of sub brands at the region-retailer-brand level. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. DC is a dummy variable taking value 1 if the supermarket is a discounter. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B8: Baseline Results: Removing Neighboring Markets

	Model 1	Model 2	Model 3	Model 4
Treat	0.0062**	0.0056*	0.0080*	0.0066
	(0.0028)	(0.0029)	(0.0041)	(0.0042)
Treat × ΔHHI		0.067*		0.12**
		(0.040)		(0.049)
Treat × *DC*			-0.0032	-0.0018
			(0.0049)	(0.0049)
Treat × *DC*× ΔHHI				-0.13**
				(0.064)
	FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls			
N	640371	640371	640371	640371
R^2	0.994	0.994	0.994	0.994

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Neighboring markets are dropped from the sample. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B9: Baseline Results: Removing Cities

	Model 1	Model 2	Model 3	Model 4
Treat	0.0037	0.0027	0.0045	0.0025
	(0.0028)	(0.0028)	(0.0038)	(0.0039)
Treat × Δ*HHI*		0.097**		0.17***
		(0.045)		(0.051)
Treat × *DC*			-0.0016	0.00047
			(0.0043)	(0.0045)
Treat × *DC* × Δ*HHI*				-0.18**
				(0.078)
N	401580	401580	401580	401580
R^2	0.995	0.995	0.995	0.995

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Cities are removed from the sample. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B10: Matching Results

Variable	Unmatched Matched	Mean Treated	Mean Control	%bias	%reduct bias	t-test t	t-test p>t
Pscore	U	.57779	.5347	42.8		10.91	0.000
	M	.56703	.56694	0.1	99.8	0.03	0.980
Δ Log(Price)	U	.01298	.01128	7.2		1.88	0.060
	M	.01302	.0128	0.9	87.0	0.32	0.750
$\Delta Log(Price)^2$	U	.00051	.00092	-16.6		-4.49	0.000
	M	.00052	.00052	-0.1	99.2	-0.08	0.933
Pop density	U	331.71	219.81	35.3		8.73	0.000
	M	278.9	277.04	0.6	98.3	0.20	0.839
Unemployment	U	.03701	.02914	5.3		1.36	0.174
	M	.03722	.03406	2.1	59.9	0.54	0.589
HH Income	U	2157	2170.1	-1.7		-0.45	0.655
	M	2158.2	2139.7	2.4	-41.4	0.66	0.512
Age	U	47.006	47.245	-2.0		-0.53	0.599
	M	46.957	47.231	-2.3	-14.7	-0.63	0.529
Children	U	.71855	.66429	6.3		1.63	0.104
	M	.72493	.76969	-5.2	17.5	-1.36	0.175

Notes: The table shows the results of a propensity score matching performed in the year 2006. The first column contains the matching variables. Δ Log(Price) is the average growth rates of prices before treatment. The other variables are pre treatment averages as well. For each matching variable there are two rows that show unmatched (U) and matched (M) means of treated and control markets. The last two columns show a t-test with H0: means of the respective matching variable of treated and control markets are equal.

Table B11: Baseline Results: Propensity Score Reweighting

	Model 1	Model 2	Model 3	Model 4
Treat	0.0038	0.0031	0.0077**	0.0064*
	(0.0024)	(0.0024)	(0.0033)	(0.0034)
Treat × Δ*HHI*		0.078**		0.14***
		(0.039)		(0.049)
Treat × *DC*			-0.0077**	-0.0065
			(0.0039)	(0.0040)
Treat × *DC* × Δ*HHI*				-0.13**
				(0.065)
	FEs for Region-Retailer-Brand, Category-Time, Retailer-Time + Regional Controls			
N	661079	661079	661079	661079
R^2	0.994	0.994	0.994	0.994

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Treatment markets receive a weight of 1 and control markets a weight of p/(1-p), where p is the probability of being treated. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. ΔHHI is the expected change in market concentration. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B12: Random Treatment Groups at Original Treatment Time

Random Draw	1	2	3	4	5
Treat	-0.00046	-0.00030	-0.00075	-0.00042	0.00028
	(0.00067)	(0.00069)	(0.00067)	(0.00065)	(0.00066)
Random Draw	6	7	8	9	10
Treat	-0.00041	0.000020	0.00041	0.00024	-0.00050
	(0.00065)	(0.00066)	(0.00067)	(0.00064)	(0.00065)
FEs for Region-Retailer-Brand, Retailer-Time, Category-Time, Time + Regional Controls					

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Treat is a dummy variable taking randomly the value 1 post merger. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Table B13: Changing Treatment Time with Original Treatment Group

Treat Time	>=2005 Q3	>=2006 Q3	>=2007 Q3
Treat	0.00099	-0.0013	-0.0018
	(0.0019)	(0.0015)	(0.0018)
FEs for Region-Retailer-Brand, Retailer-Time, Category-Time + Regional Controls			
N	475655	475655	475655
R^2	0.995	0.995	0.995

Notes: The dependent variable are log mean prices at the region-retailer-brand level. Treat is a dummy variable taking the value 1 post merger for local markets that contain both the merging parties. Post merger time varies as displayed in the first row. Regional controls consist of population density, the mean income, the mean age, the unemployment rate, and the average number of children per household. Standard errors are clustered at regional level and shown in parentheses. Significant at 1% ***, Significant at 5% **, Significant at 10% *

Chapter 3

Innovation and Institutional Ownership: Comment

Co-authored with Joel Stiebale

3.1 Introduction

In an important and widely cited paper, Aghion et al. (2013) estimate a positive causal effect of institutional ownership on innovation. They explain this effect by a career concern model, related to Holmström (1982), in which monitoring allows institutional investors to identify and reward managerial ability in risky innovation projects. The revelation of managerial ability insures good managers against unlucky innovation outcomes, which the market interprets as a negative signal of their ability, and induces them to innovate. This theory implies that institutional ownership and product market competition are complements, since competition increases the probability of imitation and hence innovation failure.[1] As argued by Aghion et al. (2013), the channels by which institutions affect innovation have important implications. If monitoring and insider ownership are major determinants of innovation, policy measures which lead to less outside board membership and higher board representation of insiders such as institutional owners would spur innovation.

This paper analyzes the impact of institutional investors on innovation and heterogeneous effects among firms which face different degrees of credit constraints and product market competition. We test the alternative hypothesis that institutional investors induce innovation by alleviating financial constraints. This hypothesis is related to a large literature on information asymmetries in capital markets which argues that suppliers of finance are confronted with an adverse selection problem leading to the rationing of finance and underinvestment (e.g. Hubbard, 1998; Stiglitz and Weiss, 1981). Research and development (R&D) is typically associated with lower collateral value but higher riskiness and asymmetric information problems compared to tangible investment. This implies that financial constraints are particularly severe for the financing of innovation (Brown et al., 2012; Hsu et al., 2014) which has been confirmed by robust empirical evidence (see, for instance, Aghion et al., 2012; Hottenrott and Peters, 2012; Stiglitz and Weiss, 1981). We argue that institutional owners may alleviate asymmetric information problems in credit markets and improve access to finance. Firms may benefit from institutional ownership directly via lower financing costs or indirectly because institutional investors' monitoring activities and financial expertise may act as a signal for creditors that their funds are used productively (see, e.g., Boucly et al., 2011).

[1] Luong et al. (2014), Bena et al. (2015) and Lee (2005) also find a positive effect of institutions on innovation. Overall, the results on institutional ownership and innovation are mixed. See Belloc (2012) for an overview of related literature.

For our empirical analysis, we use the same data set and baseline specification as Aghion et al. (2013). We extend their analysis by estimating heterogeneous effects for firms that operate in industries with high dependence on external finance. In these industries, credit constraints are particularly important since internal funds are usually insufficient to finance investment (Rajan and Zingales, 1998).[2] Our results show that the positive impact of institutional investors on innovation is concentrated in these industries. Further, we find that the complementarity between competition and institutional ownership estimated by Aghion et al. (2013) and others vanishes after financial dependence, and its interaction with institutional ownership, is controlled for. Hence, we argue that a positive interaction between institutions and competition does not stem from insurance against innovation failure but from financial dependence if this factor is omitted. We investigate the relationship between competition and dependence on external finance further by estimating separate regressions across subsamples with different degrees of competition and financial dependence.

We also find that the effects of institutional investors on innovation are concentrated among firms with initially low credit ratings, arguably firms for which financial constraints typically play an important role (Carreira and Silva, 2010; Panetta et al., 2009; Rodano et al., 2016). Further, we provide evidence that the sensitivity of R&D investment to the availability of internal funds decreases with the degree of institutional ownership. Finally, we test other empirical predictions of the career concern model, a significant impact of short-term profits on CEO turnover and a lower impact of bad performance on CEO firing with more institutional owners. We find that estimates by Aghion et al. (2013) which support these predictions only hold in industries in which financial dependence is high and thus short-term profits may be needed to finance investment or service debt. Our results are robust with respect to the model specification, the measure of financial dependence, and the application of an instrumental variable (IV) estimator.

The rest of this paper is organized as follows. Section 1 provides a description of the data, section 2 describes the econometric specification. Results of the empirical analysis are presented in section 3 and section 4 concludes.

[2] A similar empirical strategy is chosen by Boucly et al. (2011) and Amess et al. (2015) who provide evidence that buyouts undertaken by private equity firms – a specific sub-group of institutional investors – can alleviate financial constraints and thereby induce firm growth and patenting. Agca and Mozumdar (2008) find that institutional investors can reduce the sensitivity of (tangible) investment to the availability of internal funds.

3.2 Data and variables

For our analysis, we exploit a rich firm-level data set from Aghion et al. (2013) which includes 6178 observations on 800 firms.[3] It contains information on institutional ownership from Compact Disclosure, patent counts weighted by forward citations from the NBER Patent Database and accounting data including capital intensity, R&D, sales, and firm value from Compustat.

For our empirical analysis, we construct a measure of industry-level financial dependence proposed by Rajan and Zingales (1998). It proxies the desired amount of investment that cannot be financed by internal cash flow. This measure is calculated as capital expenditures minus cash flow from operations divided by capital expenditures following the variable definitions in Rajan and Zingales (1998). While financial dependence is not a direct measure of credit constraints, previous evidence indicates that financial constraints are more binding for firms in industries with high dependence on external finance. In particular, it has been shown that firms in industries with high financial dependence benefit most from stock market and banking development (Amore et al., 2013; Rajan and Zingales, 1998). Hence, if institutional investors reduce financing constraints, we expect larger effects of institutional ownership in financially dependent industries. We use data on all firms available in Compustat over the pre-sample period 1980-1990 to reduce potential endogeneity problems. Industry-level financial dependence is defined as the median of the firm-specific index for each of 39 different 3-digit SIC industries.

We also use Standard & Poor's credit ratings as an additional measure to differentiate between firms that are likely to be affected by credit constraints to a different extent. We assume that, on average, firms with low or no credit rating have to pay a higher cost premium for external funds which seems to be supported by previous research (Carreira and Silva, 2010; Panetta et al., 2009; Rodano et al., 2016). Particularly, we define firms with a rating of "A-" or higher as unlikely to be financially constrained. Table C1 in the Appendix shows summary statistics for the main variables of interests. More detailed statistics on the distribution of the credit rating variable are depicted in Table C2.

[3] The data is available at https://www.aeaweb.org/articles.php?doi=10.1257/aer.103.1.277

3.3 Empirical model

In the baseline empirical model, following Aghion et al. (2013), the conditional expectation of innovation is given by:

$$E\left(CITES_{it}|x_{it}\right) = exp\left(\alpha INSTIT_{it} + \beta \mathbf{x}_{it} + \eta_i + \tau_t\right) \tag{3.1}$$

The outcome variable, $CITES_{it}$, is computed from the number of granted patents filed by firm i in time period t. Patents are weighted by the number of forward citations to account for heterogeneity in the importance of patents. $INSTIT_{it}$ measures the proportion of equity owned by institutional investors, $\mathbf{x}_{it}$ contains control variables including sales, capital intensity, R&D stock and industry dummies, η_i is a firm fixed effect and τ_t are time dummies. Firm fixed effects are introduced into the model using the pre-sample mean of citation-weighted patents as suggested by Blundell et al. (1999).

To account for heterogeneous effects of institutional ownership on innovation, equation 3.1 is extended to allow the effect of institutional investors to vary with financial dependence:

$$E\left(CITES_{it}|x_{it}\right) = exp\left(\alpha_0 INSTIT_{it} + \alpha_1 INSTIT_{it} * FIN_{j(i)} + \beta \mathbf{x}_{it} + \eta_i + \tau_t\right) \tag{3.2}$$

$FIN_{j(i)}$ is the measure of financial dependence in industry j described in the previous section. A positive coefficient of the interaction term ($INSTIT_{it} * FIN_{j(i)}$) indicates that in industries that are more dependent on external funding, institutional investors have a larger effect. The effect of industry-level financial dependence independent of ownership is absorbed by industry dummies. We further extend equations 3.1 and 3.2 to analyze how the effect of institutional investors varies with competition as in Aghion et al. (2013) and to investigate how this effect changes when we introduce the interaction between institutional investors and financial dependence.

In an alternative specification, we interact institutional ownership with a dummy variable for firms with high credit rating. If institutional investors induce innovation by alleviating credit constraints, we should see a negative coefficient for this interaction term and a positive coefficient for $INSTIT_{it}$ since firms with high credit rating are less likely to face financial constraints. To reduce potential endogeneity problems, we use data on credit ratings from the pre-sample period (1988 to 1990).

In this specification, we include a dummy for non-ranked firms among the control variables.[4] Following Aghion et al. (2013), the main specification is estimated as a Poisson model, but we also consider alternatives including a Negative Binomial model.

3.4 Results

3.4.1 Basic results

Table 3.1 shows our baseline results. Column (1) replicates the main specification in Aghion et al. (2013) which is a Poisson model that accounts for unobserved firm heterogeneity and control variables. Institutional ownership is significantly positively associated with innovation. In columns (2) and (3), we add the interaction of institutional ownership with financial dependence. Column (3) excludes the R&D stock from the list of regressors.[5] The coefficient of institutional ownership in column (2) indicates that an increase of 1 percentage point in institutional ownership increases innovation output by about 0.37 percent when financial dependence takes a value of zero. This corresponds to an industry where the median firm's capital expenditures is equal to its operating cash flow. The interaction term between institutional ownership and financial dependence is positive and statistically significant at the 1 percent level. This suggests that in industries that are more reliant on external finance, there is a higher association between institutional investors and innovation. Starting from a situation where financial dependence takes a value of 0, an increase in financial dependence by one standard deviation raises the predicted effect of an additional percentage point of institutional ownership from approximately 0.37% to 0.6%. Columns (4) and (5) show linear regressions with the log of the number of citation-weighted patents as the dependent variable for firm-years with non-zero

[4] For 90 observations we use ratings from the sample period to exploit as much information as possible. Excluding these observations or firms with missing credit ratings from the sample did not change our results notably.

[5] Aghion et al. (2013) argue that a specification without R&D stock identifies the combined effect of institutional investors on innovation input and output while a specification that controls for R&D stock estimates the effect on innovation productivity, i.e. output conditional on innovation input. As it is not clear whether the R&D stock accurately accounts for innovation input, as indicated by the insignificant coefficients in some specifications, we prefer a broader interpretation of institutional investors on innovation. However, we believe that financing constraints are not inconsistent with an effect of institutional investors on innovation productivity since financing constraints may prevent firms from making optimal R&D investments.

patents.[6]

In column (6) and (7), we split the sample into industries with financial dependence above and below the median, respectively. In the high-dependence subsample[7], the effect of institutional ownership is more than twice as large as in the low-dependence subsample. Column (8) shows results with an interaction term between institutional ownership and $I(A)$, a dummy variable that takes value 1 for firms that are rated "A-" or higher. Following our argumentation that the impact of institutional investors is higher in more constrained firms, we would expect that they have a lower effect on firms with a high rating. This intuition is confirmed as the interaction term is negative and highly significant.

3.4.2 Institutional ownership, competition and financial dependence

Results in Aghion et al. (2013) indicate that institutional investors have a higher impact on innovation in more competitive sectors which is line with one of the predictions of the career concerns model. We argue that the main channel that drives this empirical observation is related to financial constraints. In competitive industries, firms have limited internal financial resources and have to rely more on external capital. Since the career concerns model predicts complementarity between institutional ownership and competition for reasons that are unrelated to financial constraints, we believe that it is important to control for financial dependence when this complementarity is investigated.

Table 3.2 shows results with interaction terms between institutional ownership and competition, measured as (1 - Lerner index) where Lerner index is calculated as the median gross margin at the three-digit industry level. Columns (1)-(3) replicate the results of Aghion et al. (2013) using both a time-varying measure and a time-invariant measure of competition which is computed as an average over the sample period. Columns (4)-(6) show results of analogue specifications to which we add the interaction of financial dependence and institutional ownership. Accounting for this variable, the interaction of institutional investors and competition becomes

[6] Following Aghion et al. (2013), the linear specifications contain industry fixed effects but not firm fixed effects. However, our conclusions do not change when we introduce firm fixed effects into the linear models.

[7] The median industry is assigned to the low dependence subsample and the high dependence subsample consists of all industries with larger than median financial dependence.

statistically insignificant and the parameter decreases from around 0.08 to 0.02. At the same time, the interaction of institutional ownership with financial dependence is almost unchanged compared to the results in Table 3.1 and remains statistically significant at the 5 percent level. Due to different scales, the value of the coefficients for interactions with competition and financial dependence are not directly comparable. According to the results in column (5) and descriptive statistics in Table C1, an increase in competition by one standard deviation raises the predicted effect of 1 percentage point higher institutional ownership by less than 0.06 log points. In contrast, an increase in financial dependence by one standard deviation increases the predicted effect of institutional ownership by more than 0.2 log points.

To investigate the relationship between institutional investors, financial dependence and competition further, we divide firms into four subsamples which are defined by low vs. high competition and low vs. high financial dependence. According to our financial constraints hypothesis, we expect that dependence on external finance should matter most if markups are low and hence internal finance is limited. This implies that across the two high-competition subsamples, the coefficient of institutional ownership changes more when we compare high-dependence to low-dependence industries than across the two low-competition subsamples. Similarly, we expect heterogeneity of the impact of institutions according to firms' credit rating to be more pronounced when competition is intense.

Table 3.3 shows the results of this sample split. In columns (1)-(4), we use the measure of financial dependence and in (5)-(8) we use the credit rating dummy to differentiate industries and firms according to the financial dimension. Across the low competition subsamples (columns 1 and 2), higher financial dependence is not associated with a stronger relationship between institutional investors and innovation. In contrast, financial dependence plays an important role when competition is intense. The coefficient of institutional investors is more than three times as large in column (4) where financial dependence is high than in column (3) where financial dependence is low. In columns (5)-(8), we see a similar pattern. In the low-competition subsamples (columns 5 and 6), the coefficient of institutional ownership does not differ much between firms with high and low credit ratings. In column (8), where credit ratings are low and competition is high, the coefficient of institutional ownership is large and statistically highly significantly, whereas in column (7), where credit ratings are high, it is insignificant and has a reversed sign.

3.4.3 Endogeneity of institutional ownership

It is possible that institutional investors base their investment decisions on expectations about future performance that is unobserved by the econometrician implying endogeneity of the ownership variable. To address this problem, we use the same IV for institutional ownership as Aghion et al. (2013); addition of a firm to the S&P 500 index. Aghion et al. (2013) argue that fund managers are benchmarked against the S&P 500 which induces them to invest in firms listed in this index. According to the guidelines of the S&P 500 index, it is representativeness for a firm's industry that determines addition to the index but not firm performance, innovation or investment potential.[8] The IV estimator is implemented in a control function approach where residuals from a first stage are inserted into a second stage count data regression.

Table 4.4 shows first and second stage results of the IV regression for the full sample in (1) and (2) and for industries with high and low financial dependence in columns (3)-(6). The instrument is highly significant in all first stage regressions. In columns (1) and (2), IV regressions of Aghion et al. (2013) are replicated showing a much higher coefficient for institutional ownership in the second stage compared to the baseline model. Second stage results in columns (4) and (6) confirm our previous results. The positive effect of institutional ownership is driven by firms in industries that are more dependent on external finance. The effect even becomes statistically insignificant for the low-dependence subsample.

The IV results show that accounting for endogenous selection increases the estimated coefficient of institutional ownership in the subsample with high financial dependence but not in the low-dependence subsample. Our hypothesis that institutional investors increase innovation by relaxing credit constraints is consistent with this result. Institutional investors may target companies that have high innovation potential but have been limited in their possibility to exploit these opportunities due to financial constraints. If we ignore endogeneity of institutional ownership, we therefore omit unobservables that are positively correlated with institutional ownership but negatively correlated with innovation in high-dependence industries. For the low-dependency subsample, financial constraints and hence selection on unobservables might be of lower importance.

[8] Guidelines can be found at `http://www.spindices.com/indices/equity/sp-500`. See Aghion et al. (2013) for further discussion of the IV.

3.4.4 Extensions and robustness checks

Another prediction of the career concern model is that CEO turnover becomes more likely after a decrease in profits but is less sensitive to changes in profitability in firms with institutional investors. To test this prediction, Aghion et al. (2013) regress the probability of CEO firing on lagged changes in the profit to assets ratio (profits divided by assets), a dummy variable taking value one for firms with institutional ownership above 25 percent, and an interaction term between these two variables. Columns (1) and (4) replicate the findings of Aghion et al. (2013) for two different sample periods considered in this paper. While their results are consistent with the predictions of the career concern model, they are in line with other explanations as well. For instance, a fall in profits – and hence internal funds – might be more likely to cause severe problems and lead to lay-offs when firms are financially constrained or need internal financial resources to service debt. A lower sensitivity of managerial turnover to changes in profitability in firms with institutional ownership is therefore consistent with institutional investors alleviating financing constraints.

To investigate this relationship in more detail, we analyze whether the correlations between profitability, institutional ownership and CEO turnover are more pronounced in industries with high levels of financial dependence. The marginal effects of these regressions are depicted in columns (2)-(3) and (5)-(6) of Table 3.5. The correlations found for the pooled sample seem to be entirely driven by industries with high dependence on external finance. Marginal effects are statistically insignificant for the low-dependence subsample and even reverse sign for the reduced time period which overlaps with the innovation sample (and has cleaner ownership data according to Aghion et al., 2013). This indicates that the associations between profitability, CEO turnover and institutional ownership are driven by financial factors rather than the mechanisms of the career concerns model.

As an additional tests for the importance of financial constraints, we use a more direct measure; the sensitivity of R&D to cash flow. The validity of investment-cash flow sensitivities as a measure of financing constraints has been challenged (e.g. Cummins et al., 2006; Kaplan and Zingales, 1997). However, this indicator has been applied in several recent contributions which argue that it is at least a useful measure of differences in financial constraints across different groups of firms (see, for instance, Bond and Söderbom, 2013; Brown et al., 2012; Erel et al., 2015). If institutional investors facilitate access to external finance, we expect that firms with a larger share of institutional ownership adjust their R&D to a lesser extent to increased availability of internal funds. To test this hypothesis, we estimate a dynamic model of R&D. Particularly, we relate R&D investment (scaled by a firm's capital stock) to its lagged value, institutional ownership, current and lagged values of Tobin's Q and cash flow, and interactions between institutional ownership and cash flow. We estimate the model in first differences and apply the GMM estimator proposed by Arellano and Bond (1991). As instruments, we use lagged values of all regressors except institutional ownership, a dummy variable for S&P index membership, and its interaction with lagged cash flow.

Table 3.6 shows the results for two alternative lag structures among the IVs. Column (1) treats all regressors except institutional ownership and its interaction terms as predetermined, column (2) treats all regressors as potentially endogenous. The R&D-cash flow sensitivities are significant which indicates the presence of financial constraints for R&D in the sample. While institutional ownership seems to induce higher R&D investment, it decreases the sensitivity of R&D to cash flow significantly. The estimated parameters imply that the cash flow sensitivity approaches zero if approximately 68 percent of shares are held by institutional investors.

We conducted several robustness checks which are closely related to those in Aghion et al. (2013). First, we control for firm value which might be an important, albeit potentially endogenous, determinant of innovation output (see Table C3 in the Appendix). Columns (1) to (3) of table C4 in the Appendix show results when we use 4-digit instead of 3-digit industry dummies. Results obtained from a Negative Binomial model are documented in columns (4) and (5). All these robustness checks do not change our conclusion regarding the interaction of institutional investors with financial dependence and credit ratings.

One might be concerned about the implicit assumption of a linear effect of financial dependence and the role of potential outliers. To address this concern, we rank industries according to the value of financial dependence in an alternative specification. The industry with the lowest value of financial dependence is assigned rank 1, and the industry with the highest financial dependence is assigned rank 39. Column (5) of table C4 shows that the interaction of institutional ownership with this ordinal scale of financial dependence is positive and statistically significant as well.

3.5 Conclusion

This paper builds on recent work by Aghion et al. (2013) who find a positive relationship between institutional investors and innovation. The presumed mechanism for this relationship is that monitoring by institutional investors allows them to identify and reward managerial ability in risky innovation projects. They can therefore insure managers against bad luck in the innovation process which the market might interpret as a bad signal for their ability. Our empirical analysis tests the hypothesis that institutional investors induce innovation by alleviating financial constraints.

We provide evidence that institutional investors have a higher impact in industries that are more dependent on external finance. After we control for an interaction term between financial dependence and institutional ownership, we also find that the impact of institutional investors does not significantly vary with competition which contradicts a prediction of the career concern model of Aghion et al. (2013). We argue that the previously found complementarity between institutional investors and competition is driven by financial dependence rather than a reduction of career concerns. Consistent with our argument, we find that the effect of institutional ownership on innovation is concentrated among firms with relatively low credit ratings. We also show that institutional ownership is associated with lower R&D-cash flow sensitivities.

Our results have important policy implications. Previous research has explained positive effects of institutional investors by a reduction of managers' career concerns through increased monitoring which implies that policy measures that increase board representation of blockholders relative to outsiders may spur innovation. In contrast, we argue that financial constraints are a key driver for the effects of institutional investors on innovation. We suggest that policy measures aiming to induce innovation should focus on providing firms with access to finance. Especially in industries with high dependence on external finance and low internal funds, policy measures that facilitate access to external equity by institutional investors can have a large impact on innovation activity.

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3.6 Tables

Table 3.1: Main Results with Financial Dependence

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Method	Poisson	Poisson	Poisson	OLS	OLS	Poisson	Poisson	Poisson
Dependent Variable	$CITES$	$CITES$	$CITES$	$ln(CITES)$	$ln(CITES)$	Pooled	$CITES$	$CITES$
Sample	Pooled	Pooled	Pooled	Pooled	Pooled	High Dep.	Low Dep.	Pooled
Shares of Institutions x Fin. Dep.		0.000442***	0.000458***	0.000511**	0.000408***			
		(0.000141)	(0.000141)	(0.000215)	(0.000138)			
Shares of Institutions	0.00737***	0.00369***	0.00349**	0.00350	0.00393*	0.00961***	0.00424***	0.00495***
	(0.00200)	(0.00138)	(0.00150)	(0.00247)	(0.00222)	(0.00133)	(0.000978)	(0.000781)
ln(R&D)	0.0150	0.0116			0.353***	-0.155***	0.0701	0.0155
	(0.0756)	(0.0741)			(0.0707)	(0.0422)	(0.0444)	(0.0705)
ln(K/L)	0.364*	0.368*	0.367*	0.400***	0.252***	0.651***	-0.0445	0.350*
	(0.219)	(0.220)	(0.213)	(0.130)	(0.0880)	(0.0679)	(0.0547)	(0.195)
ln(Sales)	0.149**	0.151**	0.153***	0.545***	0.275***	0.295***	0.116**	0.148
	(0.0728)	(0.0725)	(0.0255)	(0.0613)	(0.0702)	(0.0615)	(0.0548)	(0.0996)
I(A)								0.509**
								(0.207)
Shares of Institutions x I(A)								-0.0101***
								(0.00266)
N	6178	6178	6178	3998	3998	2706	3472	6178

Standard errors, clustered at the 3-digit industry level, in parentheses. * p<0.10, ** p<0.05, *** p<0.01
Column 1 replicates the results in Aghion et al. (2013).
The dependent variable is patents weighted by future citations (log patent citations in linear models).
Fin. Dep. is an industry-level measure of financial dependence.
High Dep. (Low Dep.) are industries with a value of Fin. Dep. above (below) the median.
$I(A)$ takes a value of one (zero) if firms have a rating at least (worse than) "A-".
All regressions include time- and 3-digit industry fixed effects. Firm fixed effects are used in all regressions but the linear models.
The sample period is 1991-1999

Table 3.2: Competition Interactions

	(1)	(2)	(3)	(4)	(5)	(6)
Measure of Competition	Varies	Varies	Constant	Varies	Varies	Constant
Share of institutions	0.00739***	-0.0645***	-0.0683***	0.00368***	-0.0111	-0.0133
	(0.00204)	(0.0298)	(0.0280)	(0.00138)	(0.0309)	(0.0323)
Competition	0.346	-3.691		-0.164	-0.958	
	(2.334)	(3.336)		(2.140)	(3.070)	
Share of institutions x Competition		0.0821**			0.0174	
		(0.0348)			(0.0365)	
Share of institutions x Avg. Competition			0.0868***			0.0201
			(0.0330)			(0.0385)
Share of institutions x Fin. Dep.				0.000443***	0.000392**	0.000371**
				(0.000140)	(0.000178)	(0.000188)
N	6178	6178	6178	6178	6178	6178

Standard errors, clustered at the 3-digit industry level, in parentheses. * p<0.10, ** p<0.05, *** p<0.01
The dependent variable is patents weighted by future citations.
Columns 1-3 replicate Aghion et al. (2013).
Fin. Dep. is an industry-level measure of financial dependence.
All regressions include time-, 3-digit industry-, and firm fixed effects
The sample period is 1991-1999

Table 3.3: Competition and Financial Dependency

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Competition	Low Comp.		High Comp.		Low Comp.		High Comp.	
Fin. Dependence	Low	High	Low	High	Pooled	Pooled	Pooled	Pooled
Credit Rating Dummy	All	All	All	All	= 1	= 0	= 1	= 0
Share of Institutions	0.00432***	0.00243	0.00339***	0.0106***	0.00468	0.00435***	-0.00324	0.0103***
	(0.00145)	(0.00341)	(0.00115)	(0.000139)	(0.00352)	(0.00163)	(0.00231)	(0.000970)
ln(K/L)	-0.0740	0.652***	0.0486	0.620***	-0.0766	0.164	0.426*	0.566***
	(0.0598)	(0.231)	(0.0966)	(0.0196)	(0.0547)	(0.264)	(0.237)	(0.0768)
ln(Sales)	0.132***	0.101	0.192***	0.204***	0.392***	0.0975***	-0.0931	0.233***
	(0.0415)	(0.0792)	(0.0162)	(0.00118)	(0.119)	(0.0329)	(0.118)	(0.0135)
N	1925	1189	1547	1517	545	2569	352	2712

Standard errors, clustered at the 3-digit industry level, in parentheses. * p<0.10, ** p<0.05, *** p<0.01
The dependent variable is patents weighted by future citations. All regressions include time-, 3-digit industry- and firm fixed effects.
The sample period is 1991-1999
Low and high competition (financial dependence in 1-4) is determined by the median.
Credit rating dummy takes a value of one (zero) if firms have a rating at least (worse than) "A-"

Table 3.4: IV Estimates

	(1)	(2)	(3)	(4)	(5)	(6)
Equation	1st Stage	2nd Stage	1st Stage	2nd Stage	1st Stage	2nd Stage
Method	OLS	Poisson	OLS	Poisson	OLS	Poisson
Sample	Pooled	Pooled	High Dep.	High Dep.	Low Dep.	Low Dep.
I(sp500)	8.973***		10.57***		8.010**	
	(1.990)		(1.782)		(3.231)	
Shares of Institutions		0.0258***		0.0233***		0.00494
		(0.00500)		(0.00326)		(0.0114)
ln(K/L)	-0.941	0.378*	0.182	0.603***	-1.469*	-0.0502
	(0.887)	(0.210)	(1.503)	(0.0646)	(0.786)	(0.0598)
ln(Sales)	4.557***	0.0407	4.913***	0.0817	4.277***	0.155**
	(0.677)	(0.0342)	(0.400)	(0.0538)	(0.983)	(0.0613)
1st stage residual		-0.0208***		-0.0163***		-0.00106
		(0.00506)		(0.00412)		(0.0126)
N	6178	6178	2706	2706	3472	3472

Standard errors, clustered at the 3-digit industry level, in parentheses. * p<0.10, ** p<0.05, *** p<0.01
The dependent variable is patents weighted by future citations in (2),(4) and (6).
In the first stage regressions (column 1, 3 and 5), the dependent variable is institutional ownership.
All regressions include time-, 3-digit industry-, and firm fixed effects.
Low and high financial dependence is defined as lower and larger than median financial dependence, respectively.
The sample period is 1991-1999.

Table 3.5: CEO Performance and Financial Dependence

	(1)	(2)	(3)	(4)	(5)	(6)
Sample	Pooled	Low Dep.	High Dep.	Pooled	Low Dep.	High Dep.
Time period	1988–1995	1988–1995	1988–1995	1991–1995	1991–1995	1991–1995
(Share of institutions > 25%)	1.057**	0.395	1.786**	1.364*	-0.377	3.298**
x Δ(Profits/assets)$_{t-1}$	(0.456)	(0.485)	(0.724)	(0.790)	(1.020)	(1.391)
Share of institutions > 25%	-0.0332	-0.0337	-0.0278	-0.0396	-0.00727	-0.0575
	(0.0212)	(0.0261)	(0.0299)	(0.0294)	(0.0280)	(0.0520)
Δ(Profits/assets)$_{t-1}$	-1.274***	-0.570	-2.014***	-1.668**	0.00848	-3.466***
	(0.362)	(0.398)	(0.629)	(0.690)	(0.937)	(1.268)
N	1897	961	936	1178	598	580

Standard errors, clustered at the firm-level, in parentheses. * p<0.10, ** p<0.05, *** p<0.01
The table reports marginal effects from a Probit regression.
The dependent variable takes value 1 if a manager was forced to leave and 0 otherwise.
Columns (1) and (4) replicate results in Aghion et al. (2013)

Table 3.6: Cash Flow Sensitivity

	(1)	(2)
RD_{t-1}	-0.00563	-0.0317
	(0.0872)	(0.0961)
Tobin's Q	-0.00324	-0.000604
	(0.00344)	(0.00499)
$(\text{Tobin's Q})_{t-1}$	-0.000856	-0.00406
	(0.00223)	(0.00294)
Cash l	0.257***	0.304**
	(0.0702)	(0.125)
$(\text{Cash Flow})_{t-1}$	0.290***	0.317***
	(0.0810)	(0.0842)
Cash Flow x Shares of Institutions	-0.00361***	-0.00361**
	(0.00101)	(0.00160)
$(\text{Cash Flow x Shares of Institutions})_{t-1}$	-0.00467***	-0.00560***
	(0.00122)	(0.00136)
Shares of Institutions	0.00125***	0.00114
	(0.000327)	(0.000807)
$(\text{Shares of Institutions})_{t-1}$	0.00114	0.00166***
	(0.000885)	(0.000618)
N	2749	2749
Arellano-Bond test AR(2), p-value	0.406	0.360
Hansen test, p-value	0.213	0.370
Lags as instruments	$t-1/t-3$	$t-2/t-3$

Standard errors, clustered at the firm-level, in parentheses.

* p<0.10, ** p<0.05, *** p<0.01

The dependent variable is R&D divided by the capital stock.

Both regressions contain year fixed effects.

S&P 500 and its interactions with lags of cash flow are used as instruments in both equations.

Column (1) also uses 1 to 3-year lags of cash flow, Q and lagged R&D as instruments.

Column (2) uses 2 to 3-year lags of cash flow, Q and lagged R&D as instruments.

3.7 Appendix

Table C1: Descriptive Statistics

Variable	Mean	Std.
Citation weighted patents	234.21	1016.06
Shares held by institutional investors	46.59	22.96
External funding dependence	2.82	5.27
Dummy “A-,A,A+” rated firms	0.17	0.04
Log R&D Stock	4.51	2.06
Log capital intensity	4.40	0.77
Log sales	6.49	1.92
Tobin’s Q	3.00	4.18
Dummy S&P500 inclusion	0.31	0.46
1-Lerner Index	0.86	0.03
Pre sample mean citations	4.77	2.41

Table C2: Descriptive Statistics Credit Rating

S&P Rating	Freq.	Percent	Cum.
A	318	5.15	5.15
A+	212	3.43	8.58
A-	367	5.94	14.52
B	828	13.40	27.92
B+	646	10.46	38.38
B-	772	12.50	50.87
C	503	8.14	59.02
D	220	3.56	62.58
Missing	2,312	37.42	100.00
Total	6,178	100.00	

Table C3: Controlling for Firm Value

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Dependent Variable	*CITES*	*CITES*	*CITES*	*INST*	*CITES*	*INST*	*CITES*	*INST*	*CITES*
Sample	Pooled	High Dep.	Low Dep.	Pooled	Pooled	High Dep.	High Dep.	Low Dep.	Low Dep.
Shares of Institutions	0.000356***								
x Fin. Dep.	(0.0000981)								
Shares of Institutions	0.00391***	0.00853***	0.00394***		0.0279***		0.0276***		-0.00670
	(0.00113)	(0.00110)	(0.000981)		(0.00605)		(0.00500)		(0.01000)
Tobin's Q	0.0635***	0.0670***	0.0279	0.665***	0.0487***	0.823***	0.0499***	0.390	0.0145
	(0.00740)	(0.00450)	(0.0210)	(0.169)	(0.00552)	(0.221)	(0.00128)	(0.251)	(0.0296)
S&P500				8.060***		9.444***		7.553**	
				(2.264)		(2.010)		(3.294)	
Control function	No	No	No	Yes	Yes	Yes	Yes	Yes	Yes
N	6178	2706	3472	6178	6178	2706	2706	3472	3472

Standard errors, clustered at the 3-digit industry-level, in parentheses. * p<0.10, ** p<0.05, *** p<0.01
The dependent variable is future citation weighted patents in (1)-(3),(5),(7) and (9).
Models (4), (6) and (8) are first stage regressions with institutional ownerships as dependent variable of the models (5),(7) and (9), respectively.
All regressions contain time-, 3-digit industry-, and firm fixed effects.

Table C4: Robustness Checks

	(1)	(2)	(3)	(4)	(5)	(6)
Method	Poisson	Poisson	Poisson	Negative Binomial	Negative Binomial	Poisson
Measure of Fin. Dep.	Continuous	Continuous	-	Continuous	-	Ordinal
Shares of Institutions	0.000556***	0.000451**		0.000337**		0.000157**
x Fin. Dep.	(0.000131)	(0.000189)		(0.000131)		(0.0000708)
Shares of Institutions	0.00286**	-0.0268	0.00514***	0.00370**	0.00519***	0.00204
	(0.00124)	(0.0328)	(0.000905)	(0.00152)	(0.00127)	(0.00225)
Shares of Institutions		0.0349				
x Competition		(0.0390)				
I(A)			0.438*		0.435	
			(0.243)		(0.314)	
Shares of Institutions			-0.00765**		-0.00653*	
x I(A)			(0.00357)		(0.00371)	
Industry dummies	4-digit	4-digit	4-digit	3-digit	3-digit	3-digit
N	6178	6178	6178	6178	6178	6178

Standard errors, clustered at the 3-digit industry level, in parentheses. * p<0.10, ** p<0.05, *** p<0.01
The dependent variable is future citation weighted patents. All regressions contain time- and firm fixed effects.

Chapter 4

Productivity and Foreign Institutional Ownership

4.1 Introduction

In 2013, institutional investors held assets worth over $90 trillion in OECD countries (OECD, 2013), which makes them key players in the global financial world. The 10 largest institutions manage an amount that is approximately equal to the annual GDP of the USA. Currently, institutional holdings in companies like Google and Microsoft exceed 70%. The impact of these investors in target companies has been seen critically, as they are accused of having a short-term view that might be misaligned with the long-term interests of the companies they invest in. As a response, the OECD launched a program in 2012 aimed at fostering long-term investments by institutional investors. Foreign institutional investors have especially been discussed widely in public. In an interview in 2005 Franz Müntefering, the chairman of the social democratic party at that time, referred to foreign institutional investors as locusts that destroy firm value. This statement ignited the so-called locust debate, that gained additional traction during the financial crisis. However, it is not trivial to describe an investment strategy that aims to profitably destroy firm value, as claimed by the critics of foreign institutional engagement. One might expect that actions that destroy firm value should be punished by the market in terms of share prices. In fact, economic intuition would suggest the opposite is more likely to be true. Once invested in a company, investors have incentives to increase firm value, which has also been shown empirically (Aghion et al., 2013; Bena et al., forthcoming).

The main reason why incomes differ across countries can be attributed to differences in aggregated firm productivity (Syverson, 2011). Therefore, understanding what drives productivity is of central interest. Not only does aggregate productivity differ across countries but also, conditioned on country and industry, there are large differences in firm productivity within industries and countries (Syverson, 2004).

There are various channels through which institutional investors may influence firm productivity. As shown by Aghion et al. (2013) monitoring of the management by the investors can cause managers to invest more in R&D by reducing career concerns. Also explained by a monitoring channel, Bena et al. (forthcoming) show that foreign institutional ownership fosters investments in R&D as well as in fixed and human capital. Another important channel is finance. Empirical results indicate that institutional investors impact firms by alleviating financial constraints (Boucly et al., 2011; Amess et al., 2016; Agca and Mozumdar, 2008).

Foreign institutional investors are likely to be different from domestic institutions in Europe. As summarized by Bena et al. (forthcoming), domestic institutions often have business ties with local companies (Ferreira and Matos, 2012), which implies that domestic institutions are less useful in monitoring the management (Ferreira and Matos, 2008). This can result in the management taking actions that serves their own interests, thereby hurting the company (Stulz, 2005). Furthermore, one might argue that investing in a foreign country bears additional fixed costs, relative to investing in the home country. This reasoning is similar to Melitz (2003) and would imply that only the more productive investors invest abroad, resulting in the average foreign institutional investor being more efficient than the average domestic investor. Additionally, most of the investors are from Anglo Saxon countries. Investors from these countries are more active than investors from continental Europe.[1]

I use a rich firm panel of European firms from the Amadeus database to analyze the impact of foreign institutional investors on firm productivity, measured as a residual of a production function estimation. The hypothesis is that foreign institutional investors increase firm productivity by relaxing credit constraints. Asymmetric information can lead to frictions in the credit markets (e.g., Hubbard, 1998; Stiglitz and Weiss, 1981) resulting in inefficient capital allocation and financially constrained firms. The work in this paper is complementary to Aghion et al. (2013) and Bena et al. (forthcoming) as I suggest that foreign institutional investors alleviate these constraints either directly through lower costs of capital or indirectly by signaling to creditors that financial sources are being used responsibly (Boucly et al., 2011). This means that the financial constraints channel is consistent with investors monitoring the management. To the best of my knowledge there is no work that estimates the effect of institutional investors on firm total factor productivity. Even though Bena et al. (forthcoming) find that foreign institutional investors positively impact sales and Tobin's Q, those are only crude proxies for productivity.

I find a positive correlation between foreign institutional investors on firm productivity. In particular, I find that a one percentage point increase in foreign institutional ownership increases productivity by 0.1%. Using an IV strategy, I argue that this relationship is causal. I exploit an unanticipated cash influx of foreign institutional investors due to acquisitions of firms in the investors' portfolio by firms outside the portfolio. Assuming that portfolios had been optimally chosen implies a reallocation of the newly available cash into positions the investor was holding

[1] See, for example, The Economist April 8, 2009: http://www.economist.com/node/13446602

before the acquisition.

In order to provide evidence of the financial constraints channel, I analyze wether the effect is driven by firms and industries and during a period where one would expect financial constraints to matter. At first I exploit the financial crisis as a natural experiment for credit supply reduction. Following the collapse of Lehman Brothers, there has been a drastic productivity slowdown from which we have not yet fully recovered (Duval et al., 2017; Giroud and Mueller, 2017). In particular, financially constrained firms suffered in terms of productivity and were forced to lay off workers, as shown by Duval et al. (2017) and Giroud and Mueller (2017), respectively. If institutional investors relax credit constraints, one would expect there to be a larger effect during the financial crisis, and especially in firms that are more likely to be credit constrained. Introducing heterogeneous effects for foreign institutional investors for crisis and non-crisis years shows that the average effect is mostly driven by crisis years. Then, in a next step I differentiate between firms that are more likely to be financially constrained and where foreign institutional investors should have a larger impact, if foreign institutional ownership indeed relaxes credit constraints.

A widely used measure for financial dependence has been proposed by Rajan and Zingales (1998). It proxies the desired amount of capital that cannot be supplied by internal sources within a firm at the industry level. I show that the effect of foreign institutional ownership during the crisis is larger in industries that are more reliant on external capital. I also find that the effect is driven in industries with lower markups. Ceteris paribus lower markups imply fewer internal financial resources which make a firm more reliant on external sources. Duval et al. (2017) show that firms with a larger fraction of short-term loans over assets experienced a more severe slowdown during the financial crisis. They argue that these firms were in greater need of refinancing as relatively more debt matured. I find that in industries with a larger median ratio of short-term loans, foreign institutional investors have a larger effect on productivity during the crisis. As in Duval et al. (2017), I exploit country-specific reductions in credit supply following the collapse of Lehman Brothers, measured by spreads of credit default swaps (CDS) of local banks in the respective country. Firms operating in countries with larger average CDS spreads are more likely to be credit constrained. I find that during the crisis in those countries with higher risk premiums on the CDS of banks, foreign institutional investors have a larger impact on productivity. In particular, young and smaller-sized firms are financially constrained due to a lack of collateral and credit history (Erel et al., 2013). I find that in industries that are classified as financially constrained according to industry and country, the effect of foreign institutional investors is the largest for young and small firms during the financial crisis. An alternative measure for financial constraints is how investment-sensitive firms react to cash flow shocks (Fazzari et al., 1988). I show that firms with a foreign institutional investor are less sensitive, indicating a lower degree of financial constraints.

I contribute to the literature in several ways. The work closest related to this paper is Bena et al. (forthcoming) who use a worldwide panel of listed firms and show that foreign institutional investors have a significant positive causal effect on long-term investments, such as fixed capital, R&D, and human capital that is explained by a monitoring effect of the investors. Consistent with their findings, I also show that domestic institutional investors do not affect outcome variables. This paper is different in many ways. First, with total factor productivity, I use a more general measure of firm performance as the main outcome variable. Second, exploiting the financial crisis and a variety of proxies for financial constraints, this paper tests the hypothesis that institutional investors relax credit constraints. Third, in contrast to their work, I use mostly non-listed firms for the analysis. As argued by Farre-Mensa (2016) non-listed firms are more likely to be credit constrained compared to listed firms.

Other authors also find empirical evidence that institutional investors can relax credit constraints. Boucly et al. (2011) and Amess et al. (2016) suggest that buyouts by private equity firms can cause firm growth and increase innovation activity by reducing financial constraints. Agca and Mozumdar (2008) find that institutional investors can reduce the sensitivity of investment to the availability of internal funds. Schain and Stiebale (2016) show that institutional investors increase innovation output in publicly listed firms, specifically in industries that are more dependent on external capital. Another strand of the literature exploits the unanticipated financial crisis as a shock in the supply of credit. In particular, there is Duval et al. (2017) and Giroud and Mueller (2017) that is closely related to this paper. Giroud and Mueller (2017) show that firms with a high leverage ratio decrease employment during the financial crisis relatively more compared to low-leveraged firms. Duval et al. (2017) show that productivity slows down during and after the crisis. They show that especially firms with a higher share of assets that are financed with short-term loans suffer during the crisis. Furthermore, they exploit that the crisis affected countries heterogeneously in terms of how much the credit supply declined. They use the average CDS spread of banks within a country as a proxy for how credit constrained firms are in the respective country. The authors find that firms in countries with high bank CDS spreads suffered relatively more. This indicates that financial constraints matter in terms of productivity.

This paper is different from the existing literature of financial constraints as I combine proxies for financial constraints and the financial crisis to provide evidence that institutional investors relax credit constraints. The results in this paper are complementary to the findings in Duval et al. (2017). I find that foreign institutional investors have a larger effect on productivity in firms classified as financially constrained in Duval et al. (2017) and that suffered the most during the crisis. I further show that the additional effect of foreign institutional investors during the crisis is driven entirely by firms that are more likely to be financially constrained.

The rest of the paper is structured as follows. Section 4.2 presents the data used for the analysis. In 4.3 the empirical strategy is presented. Section 4.4 shows the empirical findings. Robustness checks are conducted in section 4.4.5 and section 4.5 summarizes and concludes the findings of this paper.

4.2 Data

The main data source is the Bureau van Dijk Amadeus database.[2] It contains financial accounting data and detailed ownership information from European firms. Due to data standardization it allows for cross-country analyses. I focus on manufacturing firms from 2-digit Nace industries 10 to 17 and 20 to 32. The final estimation sample consists of 20,746 manufacturing firms from 2006 until 2013. Figures 4.1 and 4.2 show the distribution of the sample across countries and industries, respectively. 60 percent of the sample consists of Italian, Spanish, and German firms. The range is from 0.05 to 36 percent. The three largest industries in the sample are Nace codes 28, 25 and 10 that make up 38 percent of the sample. The range is from 0.15 to 13.5 percent.

For the production function estimation I use firm sales in thousand Euros as

[2]Amadeus has been used in many articles. See for instance Budd et al. (2005); Helpman et al. (2004); Konings and Vandenbussche (2005); Stiebale (2016)

the firm outcome variable, tangible fixed assets in thousand Euros to proxy capital, material costs in thousand Euros as a measure for flexible input expenditures, and number of employees for worker input. I furthermore use information on listing status and innovation activity measured by patents weighted by future citations. From the date of incorporation, I calculate firm age. Variables that are expressed in monetary values are deflated using country-specific GDP deflators from Eurostat.

Institutional ownership is defined as the total ownership in percent that includes direct and indirect holdings in a company. Institutional investors are defined as banks, insurance companies, hedge funds, mutual pension funds, private equity firms, venture capitalists, and other financial companies. Figures 4.3 and 4.4 show the distribution of foreign institutional investments by investor type across countries and industries, respectively. The distribution of foreign investors is more dispersed across countries ranging from 1 to 12 percent compared to industries with a range of 2 to 7 percent. The dominating investor types are financial companies followed by mutual pension funds. The smallest group constitutes hedge funds and venture capitalists.

Table 4.1 presents summary statistics of the main variables of interest. On average, a firm sells goods worth 85 million Euros, has a capital stock of fixed tangible assets of 15 million Euros, uses intermediate materials for 53 million Euros, employs 243 workers, and has 0.5 weighted patents.[3] Only 1.2% of the firms in the sample are listed on a stock exchange. The average holdings of foreign institutional investors are 3.67%. There are 1,646 firms that have a foreign institutional investor for at least one year in the sample. The average holdings of domestic institutional investors is 16.13%.

The data at hand is ideal for the purpose of this paper as it covers the financial crisis, contains many European countries, and has detailed ownership information for non-listed firms. This is of particular interest as non-listed firms are more likely to be financially constrained (Farre-Mensa, 2016).

In addition to the Amadeus data I use the information about merger activity from the Zephyr database that is also provided by Bureau van Dijk. This data is used to construct the instrumental variable, as is explained later. Compustat delivers the information needed for a measure of financial constraints that is constructed at the industry level and I assume that industry characteristics are similar in the US and Europe. Information about CDS spreads of European banks during the financial

[3] Innovation activity is highly clustered among a few innovative firms in the sample.

Table 4.1: Summary Statistics

Variable	mean	std
Sales	85512.15	961632
TF. Assets	15390.8	117943.1
Employees	243.7164	1873.85
Material Cost	53525.39	691061.2
For. Inst. Ownership	3.675%	17.5
Dom. Inst. Ownership	16.135%	33.3
Markup	1.58	22.25
Weighted Patents	0.58	21.99
Listed	1.2%	10.9

Notes: Sales, TF. Assets and Material Cost are given in thousand Euros. Employees is the number of workers.

crisis is taken from Ballester Miquel et al. (2013).

4.3 Empirical strategy

This section presents the empirical approach of this paper. First, I present the framework used to estimate firm-level revenue productivity. Then, I present the baseline model and the instrumental variable strategy.

I follow Ackerberg et al. (2015), henceforth ACF, to estimate firm productivity. ACF builds on Olley and Pakes (1996) and Levinsohn and Petrin (2003) and accounts for the endogeneity of production inputs. The starting point is a Cobb Douglas production function in logs:

$$q_{it} = \beta_k k_{it} + \beta_l l_{it} + \beta_m m_{it} + \omega_{it} + \epsilon_{it} \tag{4.1}$$

where q_{it} is firm revenue in logs, k_{it} denotes fixed tangible assets in logs, l_{it} denotes the number of employed workers in logs and m_{it} denotes material inputs in logs. ω_{it} is total factor productivity in logs and is unobserved by the econometrician. ϵ_{it} is an additive error term that is assumed to be exogenous with respect to the production inputs.

Material demand is a fully flexible input. This means firms decide on material expenditure in period t after observing the productivity level, workforce, and capital in period t. This implies that material demand is a function of productivity, other production inputs such as capital and labor, and foreign institutional ownership, and is given by $m = f\left(\omega_t, k_t, l_t, Inst_{t-1}\right)$. Assuming that f is monotonic in productivity, f can be inverted such that $\omega_t = f^{-1}(m_t, k_t, l_t, Inst_{t-1})$.[4]

It is assumed that productivity has the following law of motion:

$$\omega_t = g\left(\omega_{t-1}, Inst_{t-1}, P_t\right) + \xi_{it} \tag{4.2}$$

where P_t is the probability that a firm exits following Olley and Pakes (1996).[5] I assume that foreign institutional ownership influences future productivity. This is analogous to Braguinsky et al. (2015) and Stiebale and Vencappa (2016) who allow for a correlation between the acquisition of a firm by another firm and future productivity.

Inserting the expression of productivity, the production function in (4.1) can be rewritten as $q_{it} = \beta_k k_{it} + \beta_l l_{it} + \beta_m m_{it} + f^{-1}(m_t, k_t, l_t, Inst_{t-1}) + \epsilon_{it}$. Note that none of the parameters are identified as all production inputs enter f^{-1} as well. However, we get the production output net the additive error ϵ_{it} which is denoted as $\hat{\Phi}_{it}$. Productivity can then be expressed as:

$$\omega_t = \hat{\Phi}_{it} - \beta_k k_{it} - \beta_l l_{it} - \beta_m m_{it} \tag{4.3}$$

The law of motion in (4.2) is used to construct moment conditions to identify the parameters of the production function:

$$E\left[\xi_{it}\left(\beta_k, \beta_l, \beta_m\right) Z_{it}\right] = 0 \tag{4.4}$$

As proposed by Collard-Wexler and De Loecker (2016) I rely on lagged values of investments as instruments for the identification of the capital coefficient. As the authors argue, capital values in accounting data may suffer from measurement error.[6] Z_{it} further contains current and lagged values of labor and lagged values of

[4]The exact functional form of f^{-1} is unknown, however, it is approximated by a cubic polynomial of its arguments.

[5]P_t is estimated by a probit regression of exiting the market on lagged values of a cubic polynomial of k, l and m

[6]The standard procedure of using current values of capital deliver downwards biased coefficients

material. The input elasticities are estimated for each 2-digit Nace industry separately[7] using the procedure of Ackerberg et al. (2015).

Following De Loecker et al. (forthcoming) the markups can be calculated once the elasticities are estimated. The following expression is derived assuming cost-minimizing firms and is given as :

$$\mu_{it} = \beta_M \left(\alpha_{it}^M\right)^{-1} \tag{4.5}$$

where α_{it}^M denotes material expenditures over sales.

Baseline Specification

Consistent with the production function estimation framework, for the baseline specification I use a lagged dependent variable model as follows:

$$\omega_{it} = \beta_0 + \gamma_1\omega_{it-1} + \gamma_2\omega_{it-1}^2 + \gamma_3\omega_{it-1}^3 + \beta_1 Instit_{it} + \beta_2\mathbf{X}_{it} + \epsilon_{it} \tag{4.6}$$

where $\mathbf{X}_{it}$ contains 4-digit industry, country, and time fixed effects. Furthermore, I control for listing status, age, and capital intensity. As a robustness check I also include industry-country time fixed effects, to allow for industry-country-specific time trends. In order to get consistent point estimates in a lagged dependent variable model, the error term must not be auto correlated. Alternatively, I use a fixed effects estimator to check wether the results are robust with respect to the error structure.[8]

Identification

The distribution of foreign institutional investors across firms is likely to be determined by factors that are unobserved by the econometrician. Even though in model (4.6) productivity is controlled for in period $t-1$, investors are likely to base their

of capital which are close to zero.

[7]Some 2-digit industries are pooled together, as is shown in Table 4.2

[8]As described in Angrist and Pischke (2009) the FE estimator and lagged dependent variable model deliver an upper and lower bound of the point estimate, respectively. If there is unobserved heterogeneity that is firm-specific and constant over time the fixed effect estimator is the appropriate model. However, if the firm-specific heterogeneity varies over time the lagged dependent variable model is the correct choice.

investment decisions on information that is correlated with future productivity. If investors firms self-select into better-performing firms, we would overestimate the effect of foreign institutional ownership. Another reasoning might be that investors target poorly-performing firms with room for improvement.

Aghion et al. (2013) and Bena et al. (forthcoming) use inclusion in a stock index as an instrument. This approach is based on industry representativeness of the respective stock index, which creates variation in institutional ownership that is uncorrelated with the error term. Note that for the data set at hand, this strategy is not applicable, as most firms are not listed.

The identification strategy used in this paper is borrowed from Kadach and Schain (2016) which is based on cash shocks in other positions of the portfolio of an institutional investor that creates exogenous variation in the positions of the investor that were not impacted by the shock. The basic idea is that after a firm, that had initially been in the portfolio of an investor, was bought by another firm, the investor would redistribute the sudden cash influx into positions in her portfolio she had already been holding before receiving the cash. This is based on the assumption that investors choose portfolios optimally and need to rebalance portfolio weights after receiving cash.

As illustrated in Figure 4.5, consider a situation in period $t-1$ with three companies A, B, and C and an institutional investor that holds positions in companies B and C. Now, in period t company A fully acquires company B.[9] In this event the institutional investor sells her company B holdings to company A and thus experiences an influx of cash. Now, it is assumed that initially, prior to the acquisition, the portfolio of the institutional investor had been optimally chosen. Then, following Kadach and Schain (2016), a certain amount of the newly available cash is likely to be reinvested in positions that the investor was already holding before the acquisition. In our example, this would be company C which is considered a shocked company. Note that the institutional investor might also invest in another company in period t in which she did not have any holdings in. This company would not be considered a shocked company.

A shocked company is likely to experience exogenous variation due to the portfolio rebalancing of an investor following a merger, and thus the increase of institutional holdings in a shocked company identifies a local average treatment effect.

[9] Only acquisitions for 100% of the equity of the target companies are considered in order to ensure that the institutional investor actually sells her shares

Alternatively, as a robustness check, I use a propensity reweighting estimator following the approach in Guadalupe et al. (2012) or Stiebale and Wößner (2017). The authors apply this method to mergers, but it can be used analogously for the entry decision of an institutional investor. I estimate the propensity of entry by a foreign institutional investor in period t with covariates of the preceding period $t-1$. Then I use the baseline specification and a fixed effects estimator with the probability weights $1/p$ and $1/(1-p)$ for firms with and without entry of a foreign institutional investor, respectively. Additionally, I perform a one-to-one matching procedure and restrict the sample to the treated and matched control firms.[10]

4.4 Empirical results

This section presents the empirical findings. First, the results of the production function estimation are discussed in section 4.4.1. Then in 4.4.2 the baseline results are presented. Section 4.4.3 contains the results of the instrumental variable estimator. Part 4.4.4 analyzes heterogeneous effects during the financial crisis and cash flow sensitivities are estimated. Robustness checks are conducted in section 4.4.5.

4.4.1 Production function estimation

Table 4.2 shows the results of the production function estimation. The OLS estimates deliver relatively small capital and relatively large material coefficients. Firms base their decisions on current productivity that is unobserved by the econometrician. Highly productive firms have a larger demand for inputs, implying that the unobservable productivity is positively correlated with material input while at the same time a more productive firm produces more output. This results in inflated material coefficients. The ACF estimator increases all capital coefficients up to a factor of 9. At the same time, material coefficients are reduced. On average, the estimated coefficients imply a returns to scale of roughly 0.9.

Figure 4.6 shows the distribution of logged productivity demeaned at year-industry-level. The difference between the 10th and 90th percentile is 0.77. This means ceteris paribus that the 90th percentile firm can produce twice as much as the

[10] Note that these approaches identify an average treatment effect and an average treatment effect on the treated, respectively. Therefore, the estimated effects do not necessarily have to coincide with the results of the IV estimator.

10th percentile with the same production inputs. These results are consistent with Syverson (2004) who finds a dispersion of 0.65 between the 10th and 90th percentile of log TFP.

4.4.2 Baseline results

Column 1 in Table 4.3 shows an OLS regression of productivity on foreign institutional ownership controlling for time fixed effects and polynomials of lagged productivity. From column 2 until 4 country and 4-digit industry fixed effects, capital intensity, listing status, age, and citation weighted patents as well as domestic institutional ownership are added, respectively. All models show a positive, highly significant relationship between foreign institutional ownership and productivity. A one percentage point increase in foreign institutional ownership is associated with an increase of productivity by 0.027 percent. This means a 1-standard change deviation of foreign institutional ownership increases productivity by 1%.

Figure 4.7 shows the cumulative distribution of log productivity in period t for firms that have an investor in period $t-1$ and those that do not. We see that the distribution with an investor in the previous period is located almost strictly below the distribution without an investor. This means that firms with an investor in period t draw from a more favorable distribution. Figure 4.8 shows the cumulative distribution of growth rates from $t-1$ to t for firms that experienced the entry of an institutional investor in period $t-1$ and for firms that did not. Again, we see that the blue line is below the red line for more than 90 percent of the probability mass. And for the remaining 10 percent they coincide.

It is important to control for firm innovation, as Aghion et al. (2013) show a positive relationship between institutional investors and innovation and Doraszelski and Jaumandreu (2013) show a positive relationship between firm productivity and innovation activity. Not controlling for innovation could lead to biased estimates unless this is exactly the channel one aims to identify. Column 4 controls for innovation activity. There is a positive relationship between citation-weighted patents and productivity. Column 4 also controls for initial innovation activity by including pre-sample means of citation-weighted patents. The same model also shows estimates of domestic institutional investors. The point estimate is significantly negative, but economically not relevant. A 1 percentage increase in domestic ownership decreases productivity by 0.0023 percent. The negative coefficient for domestic institutions is consistent with findings discussed in the introduction, that domestic institutions likely have no or a negative impact on firm performance.

Figure 4.9 shows the same relation as Figure 4.8 but, in addition, for those firms with entry of a foreign institutional investor in period $t-1$, it distinguishes between firms that are below and above the yearly median of lagged productivity. We see that the distribution of growth rates in period t for firms below the yearly median of lagged productivity with the entry of an investor lies below the distribution of firms without entry and, furthermore, it lies mostly below the red line which is the distribution of firms with entry that were above the yearly median. Figures 4.10 and 4.11 show how this relationship plays out in the long run. They reveal that there is a lasting impact of an investor's entry only for firms with initial low productivity. In fact, for firms above the yearly median with entry the distribution of growth rates is the same as without entry, which indicates that here investors do not have a long-run impact.

4.4.3 IV results

Table 4.4 shows the IV results of the lagged dependent variable model. Columns 1 and 2 are first and second stage results of the baseline specification, respectively. Columns 3 and 4 add further control variables. First, we see that the instrumental variable is highly significant in explaining foreign institutional ownership in columns 1 and 3. Also, the Kleibergen-Paap statistic is large enough to rule out a weak instruments problem. The second stage results in columns 2 and 4 show that an increase of 1 percentage point in foreign institutional holdings increases productivity by roughly 0.1%. The point estimate is three to four times larger compared to the baseline result. This indicates that there are unobservables which have a negative impact on productivity and simultaneously cause investors to invest in the company. This is consistent with Aghion et al. (2013) and Schain and Stiebale (2016) who also find that instrumenting institutional ownership increases the point estimate. These findings are consistent with the hypothesis that institutional investors alleviate financial constraints. The unobserved financial constraints likely have a negative impact on productivity (Duval et al., 2017) and if institutional investors relax credit constraints they are likely to self-select into firms that are financially constrained, which explains why the effect is underestimated.

A concern of the identification strategy might be that firms that are considered as shocked are on a larger growth path. In this case the instrument would not be exogenous as the growth rate is a confounding factor. In order to investigate wether this is the case I regress the instrument on lagged values of productivity, a set of controls, and controlling for unobserved time constant heterogeneity. Table 4.5 shows the results of this test. All models show that there is no correlation of past productivity and shocked firms. This indicates that shocked firms are not different in terms of past productivity.

4.4.4 Financial constraints and foreign institutional ownership

This section presents the results of the impact of foreign institutional ownership during the financial crisis and in firms that are more likely to be credit constrained. First, I show that for all firms the effect is larger during the financial crisis. Then I differentiate between firms that are more likely to be credit constrained and show that the effect during the crisis is driven by those firms. Lastly, I show that firms with a foreign institutional investor are less sensitive to cash flow shocks with respect to capital investments.

Baseline financial crisis

Table 4.6 presents baseline and IV results during the financial crisis. The crisis is defined as the years 2008 to 2010. Flannery et al. (2013) use the years ranging from 2007 to 2009 as crisis years, however, as in this paper I analyze the effect of institutional investors in $t-1$ on productivity in period t, I move the window of the crisis to the years 2008 to 2010. This ensures that I estimate the effect of foreign institutional ownership that are present in the company during the crisis. The first column is the baseline specification with an interaction term of the financial crisis and foreign institutional ownership, which measures the additional effect during the financial crisis. The effect during the crisis is highly significant and almost twice as large as during non-crisis years. A one percentage point increase in foreign institutional ownership is associated with an increase in productivity of 0.038% during crisis years whereas it is only 0.02% in non-crisis years. The IV model shows that the entire effect is driven during the financial crisis. Here, a change of 1 p.p. in foreign institutional ownership increases productivity by 0.23% during the crisis and the effect is insignificant outside the crisis.

The third column splits the baseline effect into three parts. I include a pre-crisis interaction term and a linear post-crisis time trend interacted with institutional ownership, such that the coefficient of foreign institutional ownership identifies the effect in the year 2008, the sum of institutional ownership and the interaction term with the pre-crisis dummy shows the effect before the crisis while the sum of institutional ownership and the time trend multiplied by the number of years after 2008 shows the effect after 2008. We see that the effect is largest in the year 2008 and significantly declines afterward. In 2011 the effect is on the same level as before the crisis.

Dependence on external finance

Firms that have fewer internal sources to cover their investments and are thus more dependent on external finance should intuitively be more impacted by increased frictions in the credit market, as happened during the financial crisis.

A widely used measure for financial dependence was proposed by Rajan and Zingales (1998) and is constructed as the pre-sample industry median of capital expenditures minus cash flow from operations divided by capital expenditures. It measures the desired amount of investments that cannot be financed by internal sources and thus gives the proportion of capital expenditures that relies on external finance. The larger this ratio is the more likely it is that an industry is financially constrained. This is especially true during the financial crisis when credit supply is reduced. If institutional investors indeed relax credit constraints, I expect to find a larger effect during the financial crisis in industries that are more dependent on external finance. Note that even though financial dependence does not directly measure credit constraints, it positively correlates with how firms benefit from financial development (Amore et al., 2013; Rajan and Zingales, 1998).

For the empirical implementation I estimate the baseline and the IV model for industries above and below the median financial dependence separately. Table 4.7 shows the results of the sample split. We see that the interaction term of foreign institutional ownership and the crisis dummy, is only significant in the high financial dependence samples.

Industries that have, on average, lower markups should intuitively be less resilient to an exogenous credit supply reduction. The reasoning is similar to the financial dependence measure described above. Firms, that have lower markups have fewer internal financial sources to cover investments and are more reliant on external finance. Following this, we would expect a larger effect of institutional investors in industries with lower markups during the financial crisis. Equation 4.5 delivers yearly firm-specific mark ups. Then, I use the median mark up in each 4-digit Nace industry of the year 2005 to avoid endogeneity issues.

Table 4.8 shows the results of a sample split for the baseline and the IV model. As expected, there is a larger effect in industries with lower markups during the financial crisis. In the IV model the effect is solely driven in crisis years.

Rollover risk

Firms that finance a larger part of their assets with short-term credits are more likely to face problems when credit supply worsens (Custódio et al., 2013), as they, relative to firms with few short-term loans, have a larger amount of debt to mature and thus to refinance. Duval et al. (2017) use debt maturity, which is defined as current liabilities over tangibles, as a proxy for financially constrained firms and show that firms with a higher amount of short-term loans suffered more during and after the financial crisis in terms of productivity growth rates. While Duval et al. (2017) define the measure as the firm-level value in the year 2007, I worry that this approach might create an endogeneity problem with respect to institutional investors. Institutional investors that are present in a firm might influence this ratio. To avoid this, I define the measure as the pre-sample median on 4-digit industry-level. This strategy is analogous to the financial dependence measure of Rajan and Zingales (1998). Custódio et al. (2013) show that there are industry time trends and patterns for the usage of short-term debt.

Table 4.9 reports the results of a sample split of the baseline and IV model. Here, I differentiate between industries that are above and below the median industry of debt maturity. As expected, I find that firms in industries that have a high debt maturity ratio respond more to institutional investors during the crisis. In the baseline model I find that the additional effect in the crisis is entirely driven by firms with a larger rollover risk. In the IV model the point estimate of the interaction term is twice as large in this subsample.

CDS spreads

There is large heterogeneity among countries in terms of how severe the respective national banks were hit by the financial crisis following the collapse of Lehman

Brothers. Duval et al. (2017) exploit this heterogeneity to differentiate between firms that experienced larger financial frictions during the financial crisis. They use country average CDS spreads of banks demeaned at the country-level, shortly after the collapse of Lehman Brothers as a measure for financial constraints and find that firms in countries with higher spreads experienced a larger decline in growth rates of productivity.

I use the same strategy and expect that institutional investors have a larger effect during the crisis in countries with larger average CDS spreads. Table 4.10 shows the results of the sample split. In the baseline model, again the additional effect during the crisis is completely driven by firms in countries with large average CDS spreads. In the IV model the effect more than doubles compared to countries with low CDS spreads.

Small and young firms

Small and young firms are particularly likely to be financially constrained, as they usually do not have a long credit history or have little collateral to offer (Erel et al., 2013). This means one would expect a larger effect in small and young firms if institutional investors relax credit constraints. Tables 4.11 and 4.12 show results where the subsamples that are considered as financially constrained according to the proxies employed are further split up into small and young firms. Table 4.11 contains results of the subsample with high financial dependence in columns 1 to 3 and low markups in columns 4 to 6. For each proxy the baseline regression with an interaction term for the financial crisis is shown. Within this subsample columns 2 and 4 show the results for small and young firms for the two proxies, respectively. Columns 3 and 6 show the effect of the remaining firms. Table 4.12 is analogous to Table 4.11 with high rollover risk and high CDS spreads in columns 1 to 3 and 4 to 6 , respectively. The two tables show that during the financial crisis for each proxy the effect is at least twice as large for young and small firms compared to the average effect in the respective subsample.

Combining proxies for financial constraints

Table 4.13 shows the results of the above introduced proxies for financial constraints when they are applied successively. It shows the baseline specification for six different subsamples. In the first column the sample consists of firms with larger than median financial dependence and no restriction on rollover risk and CDS spreads. The second column is the same as column 1 but there are further constraints on the sample, such that it contains only industries with larger than median rollover risk. The third column then contains firms in larger than median financial dependence, rollover risk and CDS spreads. I add layers of the proxies for financial constraints on the sample, such that firms become more financially constrained from columns 1 to 3, respectively. If foreign institutional investors relax credit constraints, one would expect the effect during the crisis to get larger from columns 1 to 3. The results in Table 4.13 confirm this intuition. The interaction term increases from left to right and almost doubles in the last column compared to the first column. Figure 4.13 illustrates this relationship graphically. In both graphs the solid (dashed) line shows average TFP for firms that (do not) have an institutional investor in the previous period. The graph on the right consists of all firms that are considered to be financially constrained, as defined in column 3 of Table 4.13. The graph on the left shows all other firms. For constrained firms, starting from the year 2008, average TFP is visibly larger for firms that have an institutional investor in the previous period. On the right side the solid and dashed line proceed almost in parallel. Columns 4 to 6 show the same specifications as in columns 1 to 3, respectively, but containing only small and young firms. The effect during the crisis more than doubles for each model compared to the same model with all firms of the respective subsample. The combined effect during the crisis for the most constrained set of firms in column 6 is more than five times larger compared to the baseline regression with an interaction term for the financial crisis in Table 4.6 column 1.

Cash flow sensitivities

An alternative measure of financial constraints are investment cash flow sensitivities. The idea behind this approach is that financially constrained firms should react more to an influx of cash in terms of capital investments compared to financially unconstrained firms which can just borrow from the market if need be (Fazzari et al., 1988). Put differently, unconstrained firms optimize investments independent of the cash flow. Even though this measure has attracted criticism about how well it proxies financial constraints (Cummins et al., 2006; Kaplan and Zingales, 1997), it has found applications in recent work arguing it is at least a good proxy for differences of financial constraints (Bond and Söderbom, 2013; Brown et al., 2012; Erel et al., 2015).

If foreign institutional investors relax credit constraints, firms should show a reduced investment cash flow sensitivity. In particular, I estimate the following model:

$$I_{it} = \beta_0 + \beta_1 CF_{it-1} + \beta_2 INV_{it-1} + \beta_3 INV_{it-1} \times CF_{it-1} + \beta_4 X_{it} + \alpha_i + u_{it} \quad (4.7)$$

where I_{it-1} is the tangible fixed capital investment[11] divided by total assets, CF_{it-1} is the cash flow divided by total assets and INV_{it-1} is a dummy variable taking value 1 if there is a foreign institutional investor in the company. X_{it} contains year dummies and sales growth. If firms are sensitive to cash flow shocks and are thus financially constrained β_1 is expected to be positive. The interaction term with foreign institutional ownership is expected to be negative if financial constraints are reduced.

Table 4.14 shows the results of a fixed effects estimation. The second column additionally controls for sales growth. First, we see that financial constraints seem to be relevant in the sample. Secondly, the interaction term is negative and completely offsets the positive baseline effect, which implies that firms with an institutional investor do not adjust investments after an increase in cash flow. This indicates that firms with an institutional investor become less financially constrained.

4.4.5 Robustness checks

This section presents a series of robustness checks dealing with heterogeneous time trends, outliers, model specification, and identification strategy.

One concern might be that there are industry or country-specific time trends that attract institutional investors and the same time have a positive impact on productivity. This could be a problem in this sample, particularly as it covers the financial crisis and the European sovereign debt crisis. Both events hit the countries heterogeneously and could thus be a confounding factor. Table D1 shows the baseline results with heterogeneous time trends. Column 1 shows the baseline results with time, industry and country fixed effects. Columns 2 and 3 show the same model including industry-time fixed effects in column 2 and industry-time and country time fixed effects in column 3. Column 4 shows the baseline regression with industry-country-time fixed effects. The last model contains the same specification as in

[11]Investments are defined as the differences in tangible fixed assets plus depreciation weighted by the ratio of tangible fixed assets over total assets.

column 4 with an IV estimator. In all models the effect remains almost unchanged.

Another concern is that the results are driven by outlier firms that are extremely productive and have institutional investors on board. In Table D2 the top and bottom 1 percent of firms in terms of productivity, labor growth, sales, capital, and material costs are removed from the sample. Columns 1 and 2 show the baseline regression and the baseline IV estimator without outliers, respectively. Columns 3 and 4 are the same as columns 1 and 2 but additionally control for industry-country-time fixed effects. In all models the impact of foreign institutional investors stays the same.

As it is consistent with the estimation of the production function, I use a lagged dependent variable model as the main specification. For a robustness check I also use a fixed effects estimator. The results are shown in Table D3. The first column shows the results of a fixed effects estimator. Columns 2 to 5 are second stage regressions of IV specifications. The second column shows the results of a fixed effects IV estimator. In the third column I additionally control for institutional ownership in the period $t-2$. This mitigates the problem that the IV might be estimating a delayed effect of foreign institutional ownership of period $t-2$ on period t. Columns 4 and 5 are the same as columns 2 and 3 using a first difference estimator instead. In all specifications foreign institutional ownership is significantly positive.

The panel in the baseline specification is unbalanced. A concern might be that this drives part of the results. One might suspect that institutional investors buy companies and sell it in pieces and the original company disappears. Or firms that went bankrupt with an investor on board do not contribute to the identification of the effect and thus biases the coefficient upwards. To tackle this concern I restrict the sample to only firms that are present in all periods. Table D4 shows the baseline specification, a fixed effects estimator, and IV results for a balanced panel. In all models the effect of institutional holdings remains almost unchanged.

The identification strategy is based on a crucial assumption. An acquisition of a firm in the investors' portfolio must not be correlated with the productivity of the other firms in the portfolio. If this is violated the instrumental variable is not valid. As alternative identification strategies Table D5 shows the results of a propensity reweighting estimator in columns 1 to 3 and a one-to-one propensity score matching estimator in columns 4 to 6. Treatment in this setup is defined as the entry of a foreign institutional investor. For each identification strategy I estimate a fixed effects model and a lagged dependent variable model. As entry is modeled I include a dummy variable that takes value 1 if there is an institutional investor in the company in columns 1, 2, 4, and 5. In columns 3 and 6 I use a dummy variable that takes value 1 if entry occurs in the respective period. All models show a significant positive effect of foreign institutional ownership on productivity.

Figure 4.14 presents the result of the one-to-one matching graphically. It shows median productivity for treatment and non-treatment firms. The entry date is normalized to zero. After entry occurs the two lines visibly depart. Firms with entry of a foreign institutional investor experience a productivity growth after the entry.

4.5 Conclusion

Information asymmetries can cause frictions in the credit market, resulting in firms that are financially constrained. It has been shown that financially constrained firms suffered more during the financial crisis (Giroud and Mueller, 2017; Duval et al., 2017). This paper suggests that foreign institutional investors relax credit constraints and increase productivity. In particular, after the financial crisis foreign investors were labeled locusts and accused of destroying firm value. Contrary to this belief, it has been shown that these investors increase long-term investments (Aghion et al., 2013; Bena et al., forthcoming). Using a European firm panel containing detailed ownership information, I show that there is a positive association of foreign institutional investors and total factor productivity. I argue that this relationship is causal and apply an instrumental variable estimator. The positive impact is more pronounced during the financial crisis, which is a period where firms faced greater credit restrictions. In particular, in industries with a high debt maturity ratio, in industries with a large financial dependence on external finance, in industries with lower mark ups, and in countries with high average bank CDS spreads, firms benefit from foreign institutional investors during the financial crisis. Especially, for young and small firms, where information asymmetries are relatively more important and that can offer less collateral, the effect of foreign institutional investors are largest during the crisis. I also provide evidence that firms with foreign institutional investors are less investment-sensitive to cash-flow shocks.

The results suggest that policymakers should consider the positive financial impact on firms when evaluating foreign institutional engagement. Credit constrained firms are more likely to lay off employees in a recession, as shown by Giroud and

Mueller (2017) and if foreign institutional investors relax constraints during a recession firms may also be less likely to reduce the workforce. More general, the paper indicates that during times when information asymmetries cause frictions in the credit market, policy interventions that aim at providing credit to firms that cannot refinance in the credit market, can reduce productivity losses and thus lessen the downward pressure of an economy during a recession.

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4.6 Figures

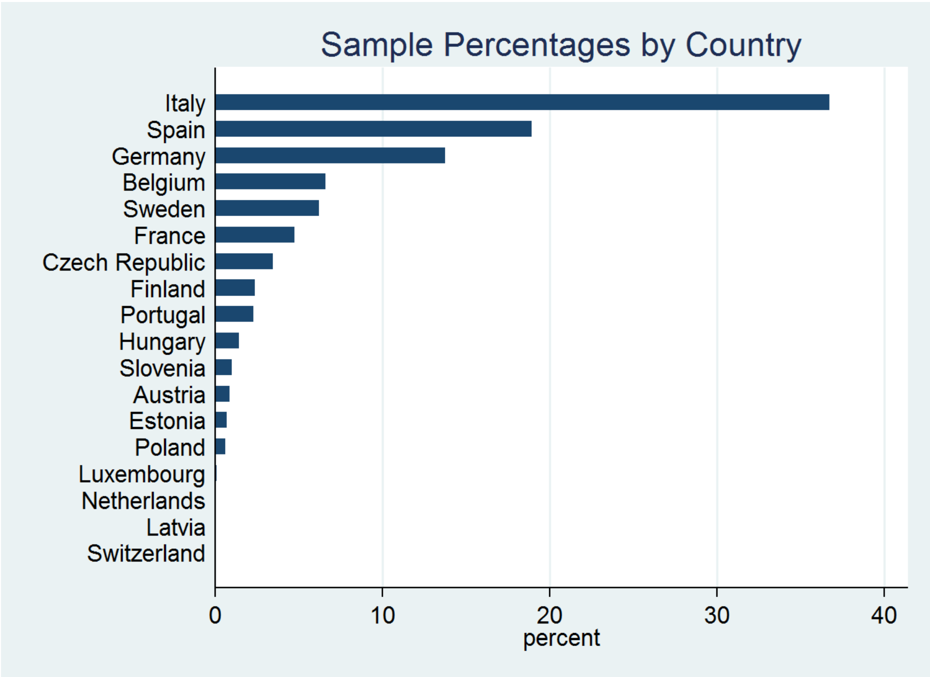

Figure 4.1: The graph shows sample percentages by country.

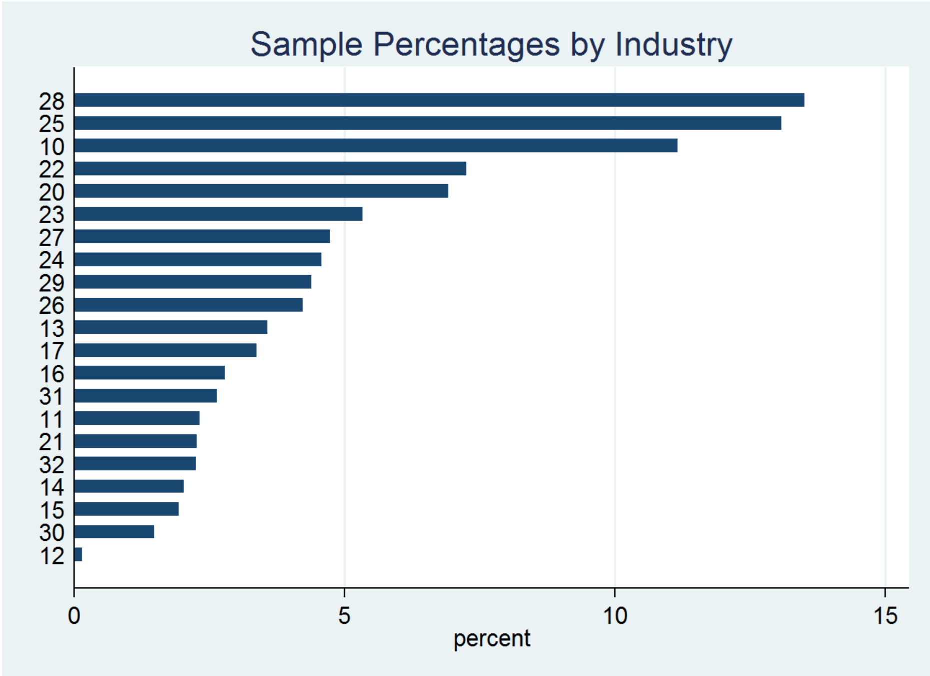

Figure 4.2: The graph shows sample percentages by Nace 2 industry.

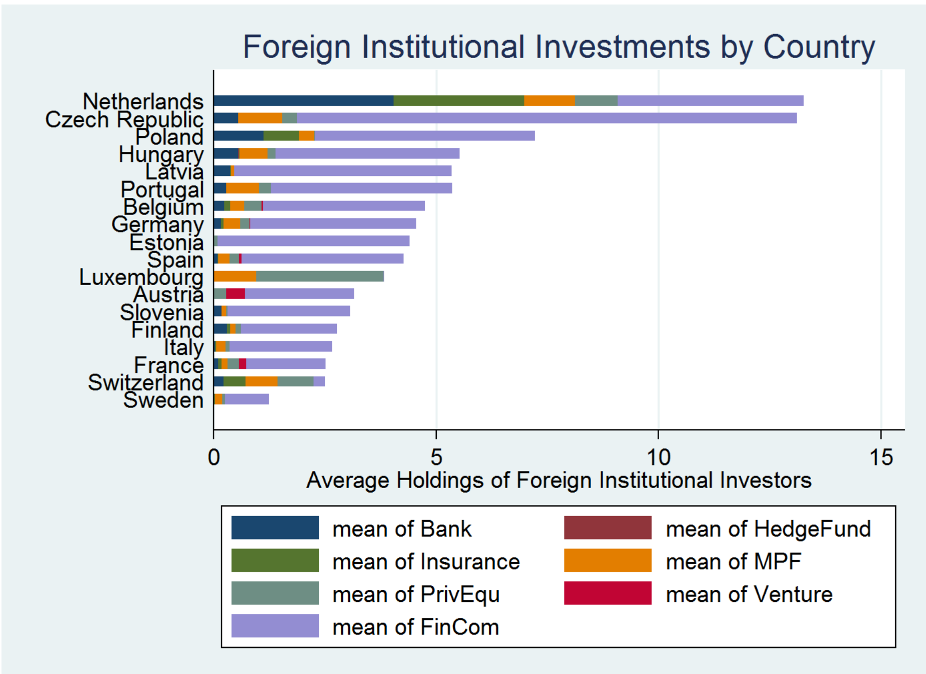

Figure 4.3: The graph shows average foreign institutional holdings by country in percent.

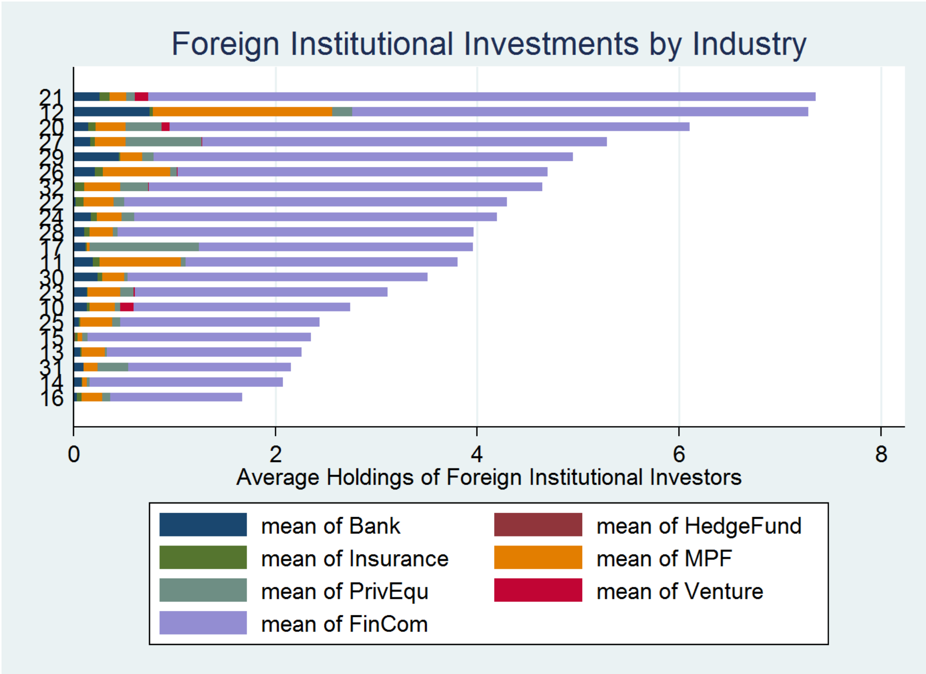

Figure 4.4: The graph shows average foreign institutional holdings by Nace 2 industry in percent.

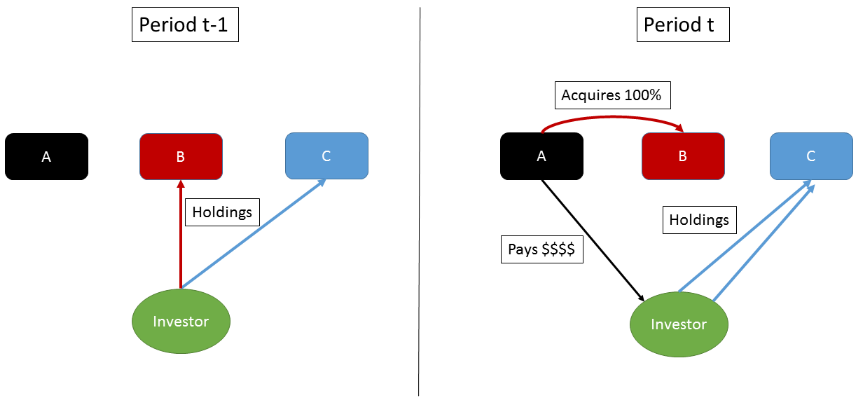

Figure 4.5: Illustration of the IV Strategy. On the left (right) hand side the situation is depicted at period t-1 (t).

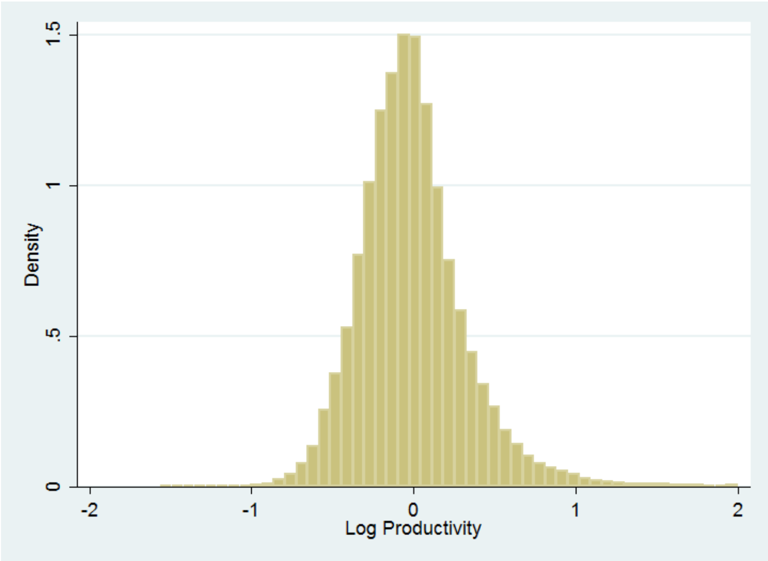

Figure 4.6: Histogram Log TFP demeaned at the country, year and 2 digit industry level

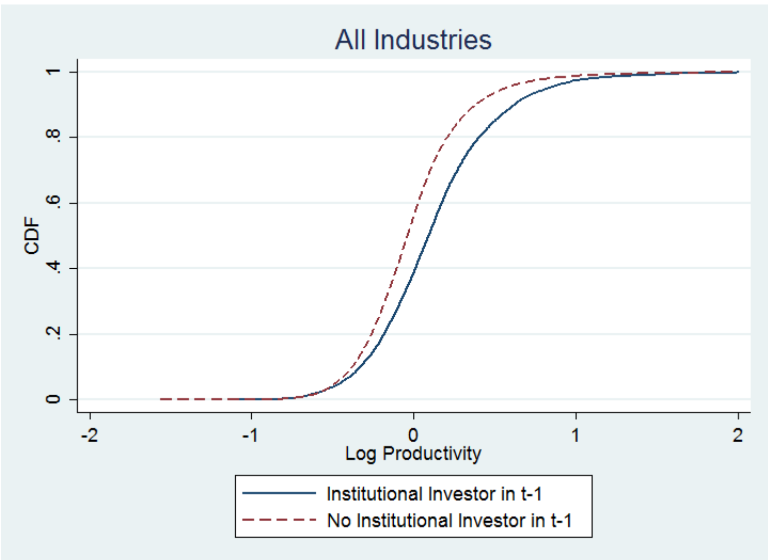

Figure 4.7: This graph show the distribution of productivity levels for firms with and without an institutional investor in the previous period

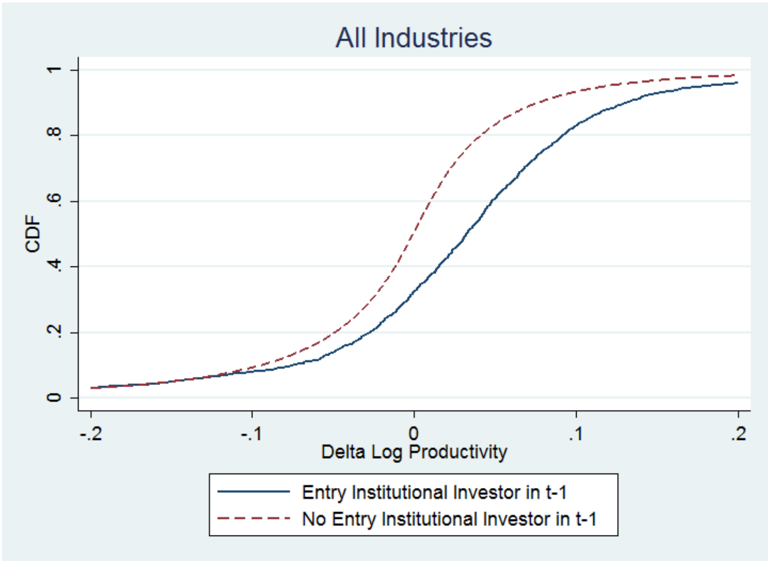

Figure 4.8: This graph shows the distribution of productivity growth rates for firms with and without entry of an institutional investor in the previous period

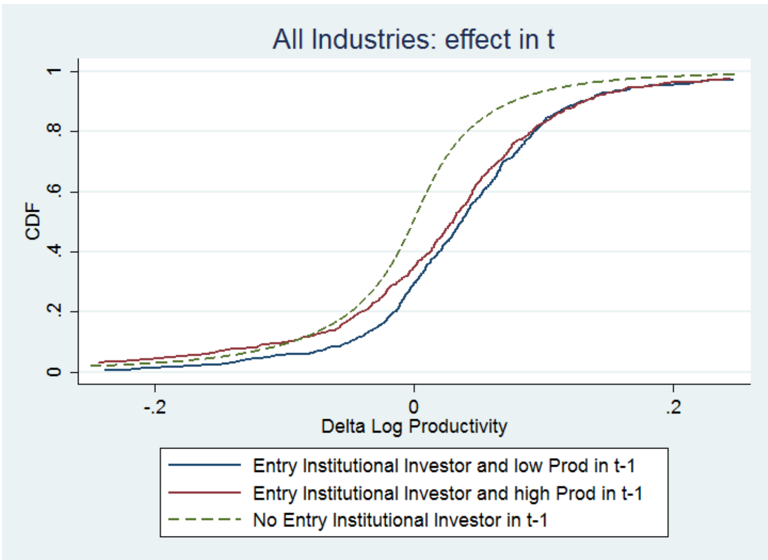

Figure 4.9: This graph shows the distribution of productivity growth rates for firms with and without entry of an institutional investor in the previous period. Additionally it differentiates between low and high productive firms in the previous period

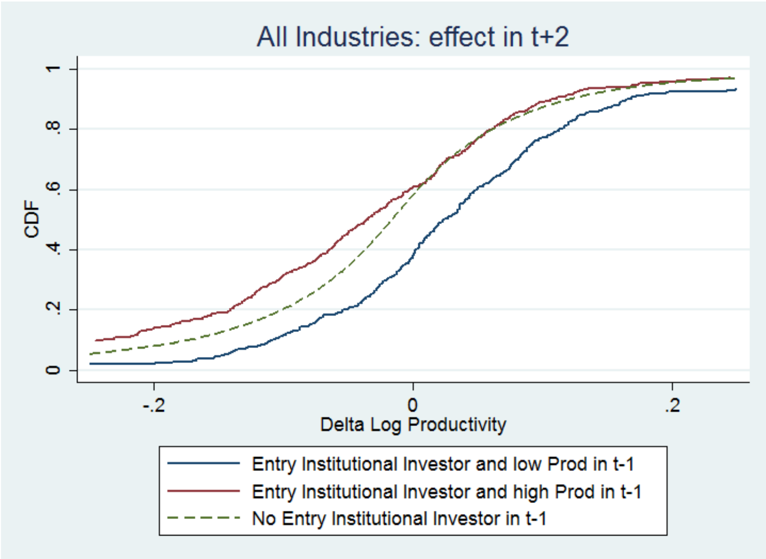

Figure 4.10: This graph shows the distribution of productivity growth rates between period t and $t+2$ for firms with and without entry of an institutional investor in the period $t-1$. Additionally it differentiates between low and high productive firms in the previous period

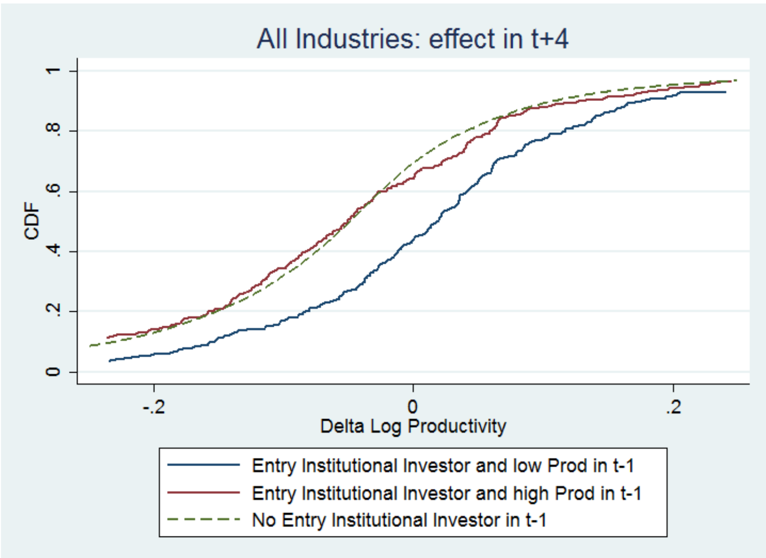

Figure 4.11: This graph shows the distribution of productivity growth rates between period t and $t+4$ for firms with and without entry of an institutional investor in the period $t-1$. Additionally it differentiates between low and high productive firms in the previous period

Figure 4.12: Yearly average log TFP in levels scaled in the year 2006 for firms with and without a foreign institutional investor in the previous period

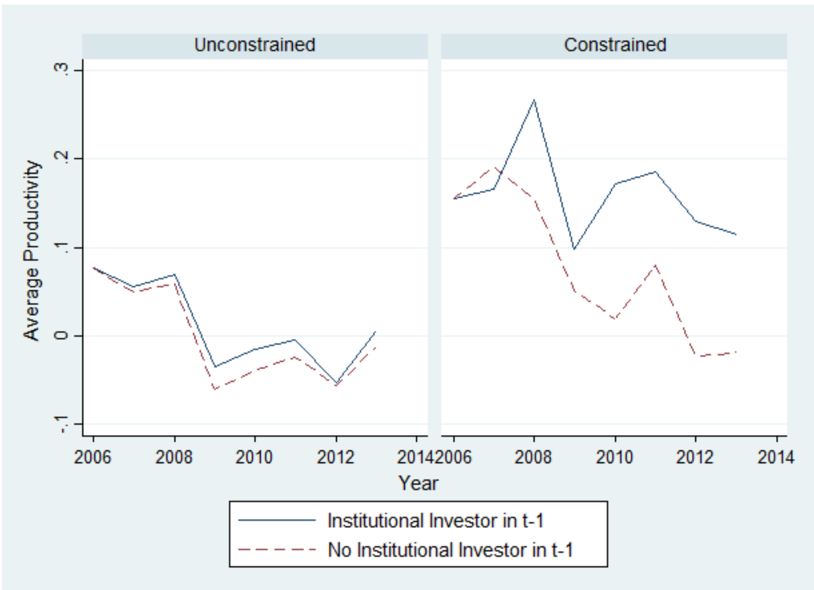

Figure 4.13: Yearly average log TFP scaled in the year 2006 differentiated between constrained and unconstrained firms. A constrained firm is defined as larger than median financial dependence, larger than median roll over risk and larger than median CDS spread.

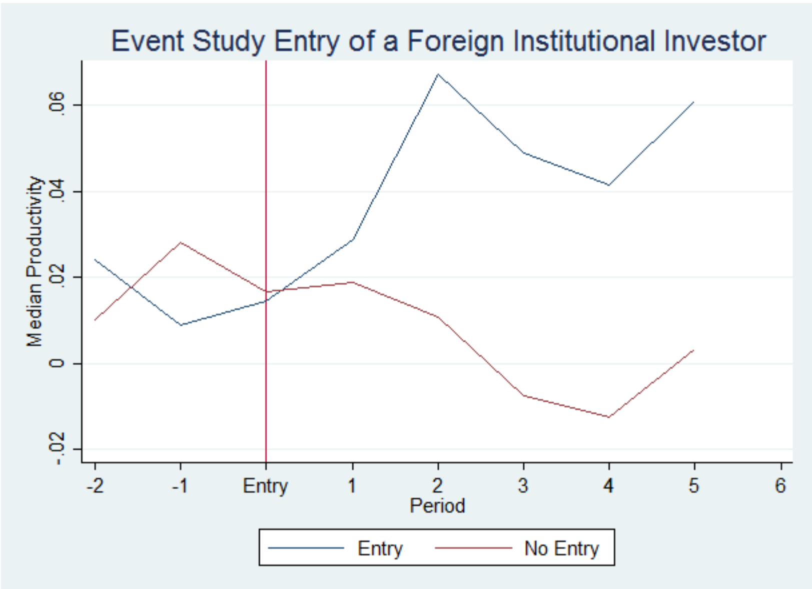

Figure 4.14: The graph shows median log productivity for firms with and without entry of a foreign institutional investor. The entry time is normalized to 0. The control firms are the result of a one to one matching procedure

4.7 Tables

Table 4.2: Production Function Elasticities

	OLS			ACF		
Industry	Cap	Lab	Mat	Cap	Lab	Mat
Food, beverages and tobacco	0.066	0.278	0.686	0.181	0.180	0.532
Textile, wearing and leather	0.019	0.258	0.639	0.076	0.247	0.434
Wood, corck and paper	0.051	0.292	0.665	0.106	0.214	0.658
Pharma and chemicals	0.062	0.339	0.599	0.281	0.289	0.272
Rubber, plastic and non-metal. min. prod	0.067	0.269	0.658	0.171	0.265	0.439
Basic and fabricated metal	0.045	0.252	0.677	0.132	0.318	0.385
Electronical products and equ.	0.023	0.309	0.629	0.211	0.304	0.336
Mach. equ. , vehicles and transp. equ.	0.018	0.299	0.655	0.163	0.218	0.525
Furniture	0.019	0.225	0.694	0.044	0.201	0.648
Other Manufacturing	0.038	0.255	0.637	0.098	0.339	0.341

Table 4.3: Baseline Foreign Institutional Ownership

Dep. Var.	ω_t	ω_t	ω_t	ω_t
For. $Instit_{t-1}$	0.0223***	0.0248***	0.0274***	0.0270***
	(0.00260)	(0.00259)	(0.00259)	(0.00260)
ω_{t-1}	0.965***	0.946***	0.939***	0.937***
	(0.00283)	(0.00307)	(0.00317)	(0.00320)
ω_{t-1}^2	0.00229	0.0143	0.0221**	0.0224**
	(0.00920)	(0.0108)	(0.0107)	(0.0106)
ω_{t-1}^3	-0.0112*	-0.0140**	-0.0161**	-0.0161**
	(0.00647)	(0.00694)	(0.00684)	(0.00681)
Log(Cap/Lab)			-0.00845***	-0.00880***
			(0.000553)	(0.000562)
Listed			0.00968***	0.00561
			(0.00375)	(0.00372)
Age			-0.0000505	-0.0000631*
			(0.0000329)	(0.0000329)
Log(1+W. Patents)				0.00568***
				(0.00141)
Log(1+ Init. W. Pat)				0.00121***
				(0.000399)
Dom. $Instit_{t-1}$				-0.00236**
				(0.00102)
N	86015	86015	86015	86015

Notes: The dependent variable is log(TFP). All models control for time fixed effects and all but the first column contain 4 digit industry and country fixed effects. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.4: Instrumental Variable Estimation

Stage	1st	2nd	1st	2nd
Dep. Var.	For. $Instit_{t-1}$	ω_t	For. $Instit_{t-1}$	ω_t
$For.Instit_{t-1}$		0.115***		0.0967***
		(0.0281)		(0.0272)
$Shocked_{t-1}$	0.162***		0.162***	
	(0.0146)		(0.0147)	
ω_{t-1}	0.0514***	0.934***	0.0511***	0.934***
	(0.00462)	(0.00370)	(0.00465)	(0.00367)
ω_{t-1}^2	-0.0281***	0.0247**	-0.0293***	0.0245**
	(0.00715)	(0.0106)	(0.00712)	(0.0106)
ω_{t-1}^3	0.00362	-0.0165**	0.00386*	-0.0164**
	(0.00231)	(0.00681)	(0.00230)	(0.00679)
Log(Cap/Lab)	0.00576***	-0.00904***	0.00559***	-0.00924***
	(0.00102)	(0.000608)	(0.00102)	(0.000609)
Listed	-0.0504***	0.00642	-0.0491***	0.00305
	(0.0100)	(0.00401)	(0.0100)	(0.00397)
Age	-0.0000727*	-0.0000446	-0.0000736*	-0.0000578*
	(0.0000406)	(0.0000334)	(0.0000409)	(0.0000333)
Log(1+W. Patents)			-0.00109	0.00568***
			(0.00293)	(0.00144)
Log(1+ Init. W. Pat)			0.000220	0.00112***
			(0.00126)	(0.000416)
Dom. $Instit_{t-1}$			-0.0330***	-0.0000650
			(0.00157)	(0.00138)
KP Stat.	123.451		122.419	
N	86015	86015	86015	86015

Notes: All models control for time, 4 digit industry and country fixed effects. $Shocked_{t-1}$ is a dummy variable taking the value 1 if the firm was shocked in the respective year. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.5: Shocked Firms and Productivity

Model	FE	FE	FE	FD	FD	FD
Dep. Var.	$Shocked_t$	$Shocked_t$	$Shocked_t$	$\Delta Shocked_t$	$\Delta Shocked_t$	$\Delta Shocked_t$
ω_{t-1}	-0.000203	-0.000100	0.00297			
	(0.00230)	(0.00235)	(0.00241)			
ω_{t-2}			-0.00413*			
			(0.00238)			
$\Delta\omega_{t-1}$				-0.00226	-0.00187	0.00219
				(0.00344)	(0.00343)	(0.00287)
$\Delta\omega_{t-2}$						-0.000119
						(0.00198)
Log(Cap/Lab)		0.00129**	0.00125*			
		(0.000620)	(0.000686)			
listed		0.274***	0.273***			
		(0.0847)	(0.0913)			
For. $Instit_{t-1}$		0.0293***	0.0222***			
		(0.00420)	(0.00438)			
$\Delta For.Instit_{t-1}$					0.0172***	0.0111***
					(0.00430)	(0.00397)
ΔLog(Cap/Lab)					0.000923	0.000885
					(0.000568)	(0.000611)
Δlisted					0.000657	0.000204
					(0.000559)	(0.000552)
N	86015	86015	57587	61974	61973	41600

Notes: All models control for time fixed effects. Columns 1-3 control for firm fixed effects. The dependent variable (Shocked) is a dummy variable taking the value 1 if a company is shocked in the current period in columns 1-3. In columns 4-6 the dependent variable is the change of the variable Shocked. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.6: Financial Crisis

Model	OLS	IV	OLS
Dep. Var.	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0205***	0.0247	0.0470***
	(0.00311)	(0.0305)	(0.00537)
$For.Instit_{t-1}$ X Crisis	0.0172***	0.209***	
	(0.00494)	(0.0583)	
$For.Instit_{t-1}$ X Trend			-0.00702***
			(0.00156)
$For.Instit_{t-1}$ X Pre Crisis			-0.0226**
			(0.0113)
ω_{t-1}	0.939***	0.935***	0.939***
	(0.00317)	(0.00368)	(0.00317)
ω_{t-1}^2	0.0221**	0.0244**	0.0221**
	(0.0107)	(0.0106)	(0.0107)
ω_{t-1}^3	-0.0161**	-0.0165**	-0.0161**
	(0.00684)	(0.00681)	(0.00684)
N	86015	86015	86015

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. The IV model shows second stage results. Crisis is a dummy taking the value 1 for the years 2008-2010. Pre Crisis is a dummy taking the value 1 for years before 2008. Trend is a variable running from the value 1 to 5 in the year 2009 to 2013 , respectively. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.7: Financial Dependence

Model	OLS	OLS	IV	IV
Fin Dep	HIGH	LOW	HIGH	LOW
Dep. Var.	ω_t	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0150***	0.0272***	-0.0388	0.120**
	(0.00454)	(0.00416)	(0.0429)	(0.0500)
$For.Instit_{t-1}$	0.0253***	0.00821	0.283***	0.120
X Crisis	(0.00722)	(0.00661)	(0.0840)	(0.0806)
ω_{t-1}	0.943***	0.933***	0.941***	0.924***
	(0.00409)	(0.00500)	(0.00452)	(0.00640)
ω_{t-1}^2	0.0208	0.0241	0.0221*	0.0297*
	(0.0130)	(0.0173)	(0.0131)	(0.0169)
ω_{t-1}^3	-0.0158*	-0.0165*	-0.0161*	-0.0174*
	(0.00896)	(0.00965)	(0.00897)	(0.00953)
N	44654	41361	44654	41361

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. The IV model shows second stage results. Crisis is a dummy taking the value 1 for the years 2008-2010. High and Low Financial Dependence are 3 digit industries with a larger and lower than median financial dependence , respectively. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.8: Mark Up

Model	OLS	OLS	IV	IV
Mark Up	LOW	HIGH	LOW	HIGH
Dep. Var.	ω_t	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0180***	0.0233***	-0.0114	0.0727*
	(0.00490)	(0.00351)	(0.0514)	(0.0375)
$For.Instit_{t-1}$	0.0226***	0.0119*	0.369***	0.0795
X Crisis	(0.00731)	(0.00635)	(0.114)	(0.0530)
ω_{t-1}	0.948***	0.922***	0.944***	0.917***
	(0.00382)	(0.00457)	(0.00477)	(0.00526)
ω_{t-1}^2	0.0118	0.0606***	0.0138	0.0648***
	(0.0105)	(0.0185)	(0.0105)	(0.0183)
ω_{t-1}^3	-0.0121*	-0.0379***	-0.0124*	-0.0386***
	(0.00657)	(0.00951)	(0.00655)	(0.00949)
N	42497	43518	42497	43518

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. The IV model shows second stage results. Crisis is a dummy taking the value 1 for the years 2008-2010. High and low mark up are industries with a larger and lower than median mark up, respectively. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.9: Roll Over Risk

Model	OLS	OLS	IV	IV
ROR	HIGH	LOW	HIGH	LOW
Dep. Var.	ω_t	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0245***	0.0172***	0.00960	0.0368
	(0.00516)	(0.00359)	(0.0624)	(0.0280)
$For.Instit_{t-1}$	0.0286***	0.00682	0.282***	0.143**
X Crisis	(0.00867)	(0.00515)	(0.100)	(0.0643)
ω_{t-1}	0.929***	0.949***	0.925***	0.945***
	(0.00383)	(0.00498)	(0.00492)	(0.00540)
ω^2_{t-1}	0.0158	0.0277*	0.0164	0.0313**
	(0.0120)	(0.0150)	(0.0121)	(0.0149)
ω^3_{t-1}	-0.0130	-0.0185**	-0.0131	-0.0191**
	(0.00905)	(0.00898)	(0.00920)	(0.00893)
N	43003	43003	43003	43003

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. The IV model shows second stage results. Crisis is a dummy taking the value 1 for the years 2008-2010. High and Low ROR are 4 digit industries with a larger and lower than median Roll Over Risk , respectively. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.10: CDS Spreads

Model	OLS	OLS	IV	IV
CDS	HIGH	LOW	HIGH	LOW
Dep. Var.	ω_t	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0162***	0.0234***	0.0478	0.0266
	(0.00540)	(0.00427)	(0.121)	(0.0393)
$For.Instit_{t-1}$	0.0260***	0.00705	0.347*	0.136**
X Crisis	(0.00874)	(0.00741)	(0.178)	(0.0656)
ω_{t-1}	0.931***	0.943***	0.925***	0.940***
	(0.00407)	(0.00428)	(0.00717)	(0.00485)
ω_{t-1}^2	0.0530***	0.00953	0.0545***	0.0113
	(0.0160)	(0.0148)	(0.0163)	(0.0148)
ω_{t-1}^3	-0.0257***	-0.00937	-0.0252***	-0.00970
	(0.00856)	(0.00807)	(0.00874)	(0.00807)
N	38021	39644	38021	39644

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. The IV model shows second stage results. Crisis is a dummy taking the value 1 for the years 2008-2010. High and Low CDS are countries with a larger and lower than median CDS Spreads during the financial crisis, respectively. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.11: Small and Young Firms I

Fin Dep	HIGH	HIGH	HIGH	ALL	ALL	ALL
Mark Up	ALL	ALL	ALL	HIGH	HIGH	HIGH
Sample	ALL	Young & Small	Old & Big	ALL	Young & Small	Old & Big
Dep. Var.	ω_t	ω_t	ω_t	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0150***	0.0132	0.0148***	0.0180***	0.0207	0.0167***
	(0.00454)	(0.0134)	(0.00469)	(0.00490)	(0.0160)	(0.00484)
$For.Instit_{t-1}$ s	0.0253***	0.0565**	0.0205***	0.0226***	0.0448*	0.0187**
X Crisis	(0.00722)	(0.0241)	(0.00743)	(0.00731)	(0.0261)	(0.00727)
ω_{t-1}	0.943***	0.905***	0.951***	0.948***	0.926***	0.954***
	(0.00409)	(0.00830)	(0.00416)	(0.00382)	(0.0101)	(0.00401)
ω_{t-1}^2	0.0208	0.0164	0.0278*	0.0118	0.0343*	0.00442
	(0.0130)	(0.0144)	(0.0154)	(0.0105)	(0.0192)	(0.0126)
ω_{t-1}^3	-0.0158*	-0.00395	-0.0231**	-0.0121*	-0.0194	-0.00982
	(0.00896)	(0.00652)	(0.00985)	(0.00657)	(0.0157)	(0.00725)
N	44654	10834	33820	42497	10661	31836

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. Crisis is a dummy taking the value 1 for the years 2008-2010. High and Low Financial Dependence are 3 digit industries with a larger and lower than median financial dependence , respectively. High and low mark up are industries with a larger and lower than median mark up, respectively. Young and Small firms are firms with an average sample age below the median and average number of employees below the median. Old and Big firms are all other firms. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.12: Small and Young Firms II

ROR	HIGH	HIGH	HIGH	ALL	ALL	ALL
CDS	ALL	ALL	ALL	HIGH	HIGH	HIGH
Sample	ALL	Young & Small	Old & Big	ALL	Young & Small	Old & Big
Dep. Var.	ω_t	ω_t	ω_t	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0245***	0.0165	0.0262***	0.0162***	0.0231	0.0137**
	(0.00516)	(0.0154)	(0.00528)	(0.00540)	(0.0143)	(0.00584)
$For.Instit_{t-1}$	0.0286***	0.0643**	0.0214**	0.0260***	0.0630**	0.0209**
X Crisis	(0.00867)	(0.0258)	(0.00906)	(0.00874)	(0.0290)	(0.00905)
ω_{t-1}	0.929***	0.916***	0.934***	0.931***	0.904***	0.937***
	(0.00383)	(0.0121)	(0.00392)	(0.00407)	(0.00959)	(0.00464)
ω^2_{t-1}	0.0158	0.0626*	0.00345	0.0530***	0.0666***	0.0509***
	(0.0120)	(0.0320)	(0.0110)	(0.0160)	(0.0240)	(0.0193)
ω^3_{t-1}	-0.0130	-0.0479*	-0.00485	-0.0257***	-0.0334**	-0.0253***
	(0.00905)	(0.0286)	(0.00612)	(0.00856)	(0.0161)	(0.00957)
N	43003	10964	32039	38021	9658	28363

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. Crisis is a dummy taking the value 1 for the years 2008-2010. High and Low ROR are 4 digit industries with a larger and lower than median Roll Over Risk , respectively. High and Low CDS are countries with a larger and lower than median CDS Spreads during the financial crisis, respectively. Young and Small firms are firms with an average sample age below the median and average number of employees below the median. Old and Big firms are all other firms. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.13: Combining Proxies for financial contraints

FIN	HIGH	HIGH	HIGH	HIGH	HIGH	HIGH
ROR	ALL	HIGH	HIGH	ALL	HIGH	HIGH
CDS	ALL	ALL	HIGH	ALL	ALL	HIGH
Sample	ALL	ALL	ALL	Young & Small	Young & Small	Young & Small
Dep. Var.	ω_t	ω_t	ω_t	ω_t	ω_t	ω_t
$For.Instit_{t-1}$	0.0245***	0.0158**	0.00911	0.0165	-0.00226	0.0171
	(0.00516)	(0.00752)	(0.0120)	(0.0154)	(0.0190)	(0.0342)
$For.Instit_{t-1}$	0.0286***	0.0435***	0.0586***	0.0643**	0.119***	0.225***
X Crisis	(0.00867)	(0.0123)	(0.0179)	(0.0258)	(0.0349)	(0.0554)
ω_{t-1}	0.929***	0.931***	0.919***	0.916***	0.908***	0.897***
	(0.00383)	(0.00479)	(0.00837)	(0.0121)	(0.0109)	(0.0130)
ω_{t-1}^2	0.0158	0.0155	0.0343	0.0626*	0.0438**	0.0724**
	(0.0120)	(0.0114)	(0.0266)	(0.0320)	(0.0189)	(0.0342)
ω_{t-1}^3	-0.0130	-0.0132*	-0.0221	-0.0479*	-0.0347***	-0.0497***
	(0.00905)	(0.00755)	(0.0140)	(0.0286)	(0.0118)	(0.0140)
N	43003	25354	11587	10964	6500	3085

Notes: All models control for time, 4 digit industry and country fixed effects and contain controls for log capital intensity, age and listing status. Crisis is a dummy taking the value 1 for the years 2008-2010. High Financial Dependence are 3 digit industries with a larger financial dependence. High ROR are 4 digit industries with a larger than median Roll Over Risk. High CDS are countries with a larger than median CDS Spreads during the financial crisis. Young and Small firms are firms with an average sample age below the median and average number of employees below the median. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table 4.14: Cash Flow Sensitivities

Model	FE	FE
Dep. Var.	$Invest_t$	$Invest_t$
$CashFlow_{t-1}$	0.0248**	0.0250***
	(0.0119)	(0.00377)
$I_{For.Inst.Inv.}$	0.0322	0.0569
	(0.152)	(0.154)
$CashFlow_{t-1}$	-0.0301***	-0.0313***
X $I_{For.Inst.Inv.}$	(0.0113)	(0.00589)
$\Delta\, Sales$		0.00309***
		(0.000400)
listed	0.447**	0.449**
	(0.224)	(0.222)
N	62194	62194

Notes: The dependent variable, cash flow and Sales are normalized by tangible assets. All models control for time. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

4.8 Appendix

Table D1: Baseline with Country Industry Time FE

Model	OLS	OLS	OLS	OLS	IV
Dep. Var.	ω_t	ω_t	ω_t	ω_t	ω_t
For. $Instit_{t-1}$	0.0274***	0.0267***	0.0260***	0.0255***	0.0971***
	(0.00259)	(0.00254)	(0.00254)	(0.00284)	(0.0308)
ω_{t-1}	0.939***	0.944***	0.944***	0.941***	0.937***
	(0.00317)	(0.00319)	(0.00319)	(0.00302)	(0.00363)
ω_{t-1}^2	0.0221**	0.0217**	0.0225**	0.0268**	0.0289**
	(0.0107)	(0.0108)	(0.0109)	(0.0115)	(0.0115)
ω_{t-1}^3	-0.0161**	-0.0164**	-0.0167**	-0.0171**	-0.0174**
	(0.00684)	(0.00697)	(0.00703)	(0.00680)	(0.00679)
Time FE	yes	no	no	no	no
Industry FE	yes	no	no	no	no
Country FE	yes	yes	no	no	no
Industry-Time FE	no	yes	yes	no	no
Country-Time FE	no	no	yes	no	no
Industry-Country-Time FE	no	no	no	yes	yes
N	86015	85942	85941	82027	82027

Notes: The dependent variable is log(TFP). The IV model shows second stage results. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table D2: Baseline and IV without Outliers and Country Industry specific Time FE

Model	OLS	IV	OLS	IV
Dep. Var.	ω_t	ω_t	ω_t	ω_t
For. $Instit_{t-1}$	0.0259***	0.0886***	0.0232***	0.0656**
	(0.00239)	(0.0266)	(0.00264)	(0.0295)
ω_{t-1}	0.963***	0.960***	0.967***	0.965***
	(0.00379)	(0.00405)	(0.00383)	(0.00411)
ω_{t-1}^2	0.0514***	0.0534***	0.0615***	0.0630***
	(0.00823)	(0.00833)	(0.00888)	(0.00898)
ω_{t-1}^3	-0.140***	-0.140***	-0.144***	-0.144***
	(0.0133)	(0.0133)	(0.0141)	(0.0141)
Time FE	yes	yes	no	no
Industry FE	yes	yes	no	no
Country FE	yes	yes	no	no
Industry-Country-Time FE	no	no	yes	yes
N	80142	80142	76182	76182

Notes: The dependent variable is log(TFP). The IV model shows second stage results. The top and bottom 1% of firms in terms of productivity, labor growth, sales, capital, and material costs are removed from the sample. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table D3: FE Estimation

Model	FE	IV FE	IV FE	IV FD	IV FD
Dep. Var.	ω_t	ω_t	ω_t	$\Delta\omega_t$	$\Delta\omega_t$
$For.Instit_{t-1}$	0.0526***	0.310***	0.223***		
	(0.00521)	(0.0842)	(0.0751)		
$\Delta For.Instit_{t-1}$				0.332***	0.218***
				(0.107)	(0.0562)
$I_{For.Instit.\,t-2}$			-0.0116**		0.0227***
			(0.00574)		(0.00829)
$I_{For.Instit.\,t-1}$					
lkl	-0.0685***	-0.0684***	-0.0669***		
	(0.00339)	(0.00340)	(0.00373)		
listed	0.0263	0.0234	0.0400		
	(0.0315)	(0.0327)	(0.0373)		
Δlkl				-0.0619***	-0.0618***
				(0.00414)	(0.00414)
Δlisted				0.0273	0.0231
				(0.0318)	(0.0283)
N	86015	86015	78514	78918	78918

Notes: All models control for time fixed effects. The IV models show second stage estimates. Standard errors are clustered at firm level and are shown in parentheses. $I_{For.Instit.}$ is a dummy variable taking the value 1 if there is a foreign institutional investor in the company. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table D4: Results for Balanced Panel

Model	OLS	IV	FE	FE IV
Dep. Var	ω_t	ω_t	ω_t	ω_t
For. $Instit_{t-1}$	0.0290***	0.100**	0.0571***	0.299***
	(0.00361)	(0.0484)	(0.00699)	(0.0986)
ω_{t-1}	0.944***	0.940***		
	(0.00341)	(0.00434)		
ω_{t-1}^2	0.0124	0.0149		
	(0.0125)	(0.0125)		
ω_{t-1}^3	-0.0117	-0.0121		
	(0.00823)	(0.00821)		
N	52336	52336	52336	52336

Notes: All models control for time fixed effects. Columns 1 and 2 control for industry and country fixed effects. The IV models show second stage results. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Table D5: Alternative Identification Strategy

Model	FE	OLS	OLS	FE	OLS	OLS
Dep Var	ω_t	ω_t	ω_t	ω_t	ω_t	ω_t
Estimator	Propensity Reweighting			One to One Matching		
$I_{For.Instit.\,t-1}$	0.0386***	0.0285***		0.0425***	0.0269***	
	(0.00541)	(0.00386)		(0.00407)	(0.00279)	
For. Instit. Entry			0.0333***			0.0405***
			(0.00540)			(0.00318)
ω_{t-1}		0.905***	0.915***		0.946***	0.950***
		(0.00898)	(0.00858)		(0.00790)	(0.00780)
ω_{t-1}^2		0.0215	0.0148		0.0638***	0.0610***
		(0.0205)	(0.0204)		(0.0185)	(0.0184)
ω_{t-1}^3		-0.0104	-0.00941		-0.0437***	-0.0433***
		(0.00841)	(0.00840)		(0.0125)	(0.0125)
N	81085	81085	81085	10566	10566	10566

Notes: The dependent variable are log(TFP). All models control for time fixed effects. The OLS models also contain 4 digit industry and country fixed effects. The propensity reweighting models contain sample weights 1/p and 1/(1-p) for treat and control group, respectively. $I_{For.Instit.}$ is a dummy variable taking the value 1 if there is a foreign institutional investor in the company. For. Instit. Entry is a dummy variable that is equal to 1 if the firm experienced entry of a foreign institutional investor in the previous period. Standard errors are clustered at firm level and are shown in parentheses. Significant at 1% $***$, Significant at 5% $**$, Significant at 10% $*$

Chapter 5

Conclusion

The ownership structure of firms is a major factor that determines how well a single firm and an entire industry performs. Ownership concentration can lead to price distortions and allocation inefficiencies or improved performance if the synergy gains of common ownership are large enough. Antitrust agencies have to take both forces into account when deciding on a merger. This thesis proposes a framework in which the potential synergy gains of hypothetical mergers can be estimated ex ante and be incorporated in a merger simulation model to estimate the likely price effects, taking into account positive and negative effects on prices. Implementing the methodology in the French dairy dessert industry using household panel scan data from 2011, we estimate that approximately half of the potential mergers are likely to generate efficiency gains. The average cost reduction is 2.5%, showing that the ad-hoc rule of 5% is too large. Depending on the size of the efficiency gains mergers can then be pro-competitive.

There is only scarce empirical evidence of synergy gains post-merger. Theoretical predictions in the retail industry are particularly ambiguous. This thesis analyzes a merger in the German retail industry using a difference-in-differences estimator exploiting variation of prices in local markets and over time. We find that, on average, prices increase post-merger. Especially, in markets with a high expected change in concentration, prices increase post-merger. A novel identification strategy allows us to identify negative price effects post mergers. We provide evidence that, in particular, in highly competitive markets synergy effects are passed through to the consumers. We also show that the remedies imposed were not sufficient to avoid anti-competitive effects.

This thesis provides evidence that efficiency gains actually can offset the positive market power effect. In the retail industry pre-merger market shares of the merging parties can be used as a predictor of likely price changes post-merger. More evidence in other industries on the presence of efficiency gains are needed. Especially the identification of the sources and what firm characteristics and which market structures are more likely to produce efficiency gains post-merger. In the retail industry there are two main sources of efficiency gains. On the one hand there are logistical savings as the merging retailers can benefit from the warehouse locations of the other party. On the other hand, the increased bargaining power may cause lower wholesale prices that can be passed through to consumers.

Frictions in the credit market caused by information asymmetries can lead to under-supply of credit and financially constrained firms. In particular, these firms are impacted by credit contractions during economic downturns. There is empirical evidence that institutional investors can alleviate financial constraints and increase long-term investments, which contrasts the public opinion that has a controversial view on institutional investors, accusing them of having too short-sighted an agenda. This thesis shows a positive association between institutional investors on innovation activity that is more pronounced in industries that are more likely to be credit constrained in the US during the 1990s. It is argued that the relationship is causal, exploiting random variation in institutional holdings following the addition of the firm into the S&P500 index. We also provide evidence that firms become less sensitive to cash flow shocks with respect to R&D investments when there is an institutional investor present in the firm.

Foreign institutional investors are likely to be different from domestic institutions as they are less likely to have business ties with the local firms and thus better act as a monitor. This thesis estimates a positive effect of foreign institutional ownership on firm productivity. I argue that the relationship is causal by using an instrumental variable estimator that exploits a sudden cash influx of investors caused by an acquisition of firms in the portfolio of the investor. The investor then proceeds to reinvest the cash into positions she was holding before the acquisition. I show that the impact is more pronounced during the financial crisis and in industries that are more likely to be credit constrained. Particularly young and small firms that lack credit history and collateral benefit from foreign institutional investors during the crisis.

Regulators and policymakers should take into account financial aspects of foreign institutional engagement. In recent contributions it has been shown that particularly financially constrained firms suffered during the financial crisis. This thesis suggests that institutional investors can reduce productivity slowdowns in firms relaxing the the financial situation. Aggregated, this could potentially lead to a more robust economy in economic downturns.

There is large evidence that institutional investors can alleviate financial constraints. For future research it is interesting to identify what kind of investors actually produce the empirical evidence. Also, it remains unclear how exactly institutional investors provide more liquidity to the firms. It is argued that the monitoring channel is consistent with the financial constraints channel as better monitoring signals to the market that funds are being used responsibly.