Firm types and heterogeneous consumption-tax incidence

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Abstract

Using very detailed micro data from VAT reforms in Sweden and Finland we show that the tax incidence of these reforms depend heavily on underlying firm-level characteristics of the price-setting firms. Our data are drawn from the restaurant industry where independent restaurants are competing, side-by-side, with restaurants belonging to chains or franchises. We show that the immediate impact of the reforms on actual prices is virtually non-existent for small independent businesses. In contrast, we find a full pass-through for a large fraction of the restaurants that belong to chains. The standard analysis of tax-incidence is heavily focused on the elasticities of supply and demand and the degree of competition, but our analysis shows that the price responses differ dramatically between independent businesses and chains, despite the fact that they operate in the same well-defined markets. Furthermore, we show that the difference in behavior between independent businesses and chains remain if we account for location, initial prices and other indicators of market segment. Notably, the zero pass-through we find for independent firms can only be reconciled with standard tax-incidence models if demand is infinitely elastic, whereas the perfect pass-through we find for many chains would predict the opposite. As these are highly unlikely scenarios, the results appear to imply that price setters at independent businesses chose to completely ignore the (highly publicized) VAT-reforms, suggesting that their price setting strategies (or objective functions) aim for (very) crude price ranges rather than fine-tuned optimized prices. In line with this hypothesis, we show that independent businesses to a much larger extent than other businesses rely on round-number pricing.

Keywords: firm heterogeneity, VAT, price incidence, restaurants

JEL-codes: [H22, H32, E31]

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1 Introduction

An increasingly active literature within public finance explores the incidence of consumption taxes (Carbonnier 2007, Doyle and Samphantharak 2008, Marion and Muehlegger 2011, Kopczuk et al. 2013, Kosonen 2013). This literature has asserted that tax incidence depends on market level conditions, such as elasticities of demand and supply and the degree of competition among firms (e.g. Myles 1989 and Weyl and Fabinger, 2013). A maintained assumption is, however, that all firms are identical, or only differ in the market conditions they face. In this paper, we relax this assumption and document substantial heterogeneity in the way different types of businesses react to the same VAT reform.

Our analysis studies the extent to which different types of businesses respond to changes in consumption taxes. Empirically, we analyze price responses to VAT-reductions in the restaurant industry in Finland and Sweden using very detailed micro data on the anatomy of price changes while separating between independent restaurants and restaurants that belong to chains or franchises.

The existing literature on tax incidence has asserted that price responses to consumption taxes depend on market-level factors. In the case of perfect competition, the price response therefore only depends on the elasticities of supply and demand. If firms instead have some price-setting power, the incidence will also depend on the exact form of imperfect competition, and on the nature of interactions between firms. But a key assumption imposed across the wide class of models used in the analysis of tax incidence is that the incidence is deterministic as long as the market is sufficiently well-defined. As a consequence, tax incidence for similar products sold in the same markets should be independent of the internal characteristics of the selling firms.

We use data on VAT-reductions in Finland during July 2010 and in Sweden January 2012 and rely on the price evolution in neighboring countries to control for time trends. To facilitate the analysis, we collected data on the anatomy of price changes across time for a representative sample of restaurants of different types in the relevant countries and matched these to administrative records held by the tax authorities in the two countries. These data allow us to dig much deeper into the price responses than would have been possible with standard CPI-type data which are much less detailed. In particular, we are able to follow the prices of the same meals over time which allows us to examine the full distribution of price changes for different types of firms. We are also able to provide a detailed account of the economic environment faced by the firms.

We divide the restaurants into two categories; we refer to restaurants that do not belong to chains or franchises as *Independent* and other restaurants as *Chains*. This dichotomy is based on the notion that pricing decisions in independent restaurants are likely to be made by owners or other on-site managers who are responsible for multiple decisions, including many of a very practical nature such as staffing and cooking, whereas pricing decision in chains are more likely to be made by specialists. Although we expect the dichotomy of firm-types to exist throughout the economy with variations both within and across industries, we believe that the restaurant sector is very well-suited for an analysis of firm-side heterogeneity since small (sometimes family-run) businesses are competing side-by-side with chains and franchises.[1] We show that Independent restaurants and Chains operate in similar market segments as both groups feature fast-food restaurants serving hamburgers and pizzas, as well as finer restaurants serving expensive dinners. As a consequence, the price distributions of independent restaurants and chains show a very substantial overlap.

The results show strikingly clean price change patterns. The overall pass-through is fairly low, a quarter of full pass-through in the short run. However, this average pass-through masks considerable heterogeneity. The immediate impact of the reforms on actual prices is virtually zero for the group of independent restaurants, whereas a large fraction of the chain restaurants instead show a full pass-through. We perform a number of robustness checks to see if this difference is due to the location, initial price (an indicator of market segment) or type of restaurant and find no support for that notion. Most notably, we show that the difference remains throughout the initial price distribution and when only analyzing establishments located close to each other within the same mall.

Our most striking finding is the immediate pass-through of zero for independent businesses. Combined with apparent price-setting power, this result can only be explained by standard tax incidence models if demand is infinitely elastic which seems unlikely, in particular since many of the chain restaurants responded so heavily. Our hypothesis is instead that the price setters at independent restaurants chose to ignore the reforms because they rely on simplified pricing strategies, i.e. they aim for cruder price targets and therefore do not re-optimize in response to the VAT reductions. This could either be due to an inability to process the (highly publicized) information regarding the reform, or due to non-standard objective functions.[2] It should

[1]On the other hand, it may be argued that the choice of industry may have geared the analysis towards finding results that are more in line with standard theory since the restaurant industry may be closer to perfect competition than many other industries. However, our results are quite far from the predictions of perfect competition models.

[2]In this context, it is also notable that the month-to-month and year-to-year variances in taxed turnover within firms is enormous. This suggests that firms face large idiosyncratic shocks and therefore they may find it difficult to learn about their demand by experimenting with small changes in prices.

be noted that the basic idea that some businesses may be behaving in ways that differ from standard theory, although absent in the literature on tax incidence, has received empirical support within other strands of the literature. Experimental evidence in Bloom and Van Reenen (2010) suggest that the quality of managerial practices vary widely across firms. Lazear (2004, 2005) asserted that smaller entrepreneurial firms tend to be run by generalists who need to attend to multiple, sometimes complicated, tasks. Drexler et al. (2014) found that entrepreneurs benefit more from simplified rule-of thumb management training than standard training. In addition, it seems likely that the behavior of price setters within independent businesses could resemble the behavior of regular consumers (at least more than professional price setters) and these have been shown to be affected by "non-standard" elements such as passive savings decisions and tax salience (Chetty et al. 2009, Finkelstein 2009 and Chetty et a. 2014).

To validate the hypothesis that independent businesses rely on cruder pricing strategies, we have analyze two indicators of price setting strategies during normal times. We first analyzed the frequency of price changes in the absence of reforms, and show that independent restaurants change their prices less frequently. We then analyzed the types of prices the restaurants set (again, in normal times) and show that independent restaurants are considerably more likely to use prices that are rounded to integer values on their meals (even after accounting for market factors, such as the price range).

It can also be noted that although the tax incidence pattern for larger firms is more in line with standard models *on average*, their distributions of pricing choices do not fit the standard models perfectly either. Somewhat simplified, our results suggest that chains either responded by fully shifting the reduced VAT to prices or by not changing their prices at all. The high frequency of full pass-through (which partly was reverted in the long run) is more consistent with highly strategic pricing behavior, where the large price reduction in the short term was intended to either receive good-will from customers, or served as a part of a coordinated effort to prevent the governments from resetting the VAT rates. We show further evidence of strategic behavior by analyzing data from the introduction of the Euro in

Estonia. According to standard theory a currency change should not change the price level, but empirically they have been found to do so (Cavallo et al. 2014) and one possible explanation for this result is that the reduced visibility of price changes during the transition can be used as a way to reduce the visibility of price increases. We find that the chains, but not independent restaurants, appear to have used the introduction of the Euro as an opportunity to increase their prices more than otherwise. This suggests that chains (but not independents) utilized the currency change as a strategic opportunity to change their prices.

Our last piece of evidence is the price dynamics after the VAT reforms. We followed the prices of the same meals for up to 18 months after the reforms. Interestingly, the restaurants belonging to chains that initially responded with a full pass-through started to increase their prices again already three to six months after the reform. Instead, those firms that did not initially reduce their price at all (effectively ignored the reform), kept their prices constant more often throughout our surveillance period. This evidence highlights that the full pass-through is unlikely to occur because of inelastic demand for some restaurants. Moreover, this observation could explain differences in price dynamics between the Finnish and Swedish reforms. Finnish prices decreased more, initially, and then gradually caught up with the control group over time. Instead, Swedish restaurants reduced their prices had a relatively modest short-run impact, but started to divert from the counterfactual price development over time. Our representative survey data shows that there are more restaurants belonging to chains or franchises in Finland than in Sweden.

Overall, we believe that our results provide strong evidence for the notion that tax incidence, in the short and long run, depend on the structure of the firms populating the market. This is in strong contrast to the assumption that market factors alone are important for the predicting the price responses to changes in VAT rates. In particular, our results suggest that independent businesses (at least in the restaurant industry) rely on crude pricing rules that reduces the frequency of price changes and increases the use of round number prices, and that this practice has real consequences for the short and medium incidence of consumption tax reforms.

The paper is structured as follows. Section 2 presents standard tehory applied to tax incidence in the literature. Section 3 presents the institutions related to the VAT reforms and explains the mechanisms we think are behind the empirical observations. Section 3 describes the data and methods. Section 4 first shows descriptive evidence on the evolution of prices for the two types of firms during the two reforms and then proceeds to estimating regressions using external control groups to identify the causal average effects. Lastly, section 5 concludes.

2 Predictions from standard tax incidence models

In this short section we highlight a few very basic results from standard tax incidence models that are useful as a background when interpreting our analysis and results.

A key result arising from the economic theory is that the effects of taxes on prices and other outcomes depend on how markets work. In the simplest, perfect competition case, we assume market clearing and price-taking behavior so that demand (D) equals supply (S) in equilibrium in a world where both are functions of prices (p). If we introduce a tax t we get $D(p) = S(p-t)$ and the standard tax incidence formula:

$$\frac{dp}{dt} = \frac{\epsilon_S}{\epsilon_S + \epsilon_D} = \frac{1}{1 + \frac{\epsilon_D}{\epsilon_S}}$$

where $-\epsilon_D$ is the demand elasticity and ϵ_S the supply elasticity. Thus, the elasticities of demand and supply are the sole determinants of tax incidence and the more inelastic side bears the burden of taxation. This implies, e.g., that to explain a zero pass-through, one needs to assume perfectly elastic demand or perfectly inelastic supply, and to get a full pass through, conversely the demand elasticity instead needs to be zero or supply elasticity infinite. Starting from non-zero tax levels, or assuming ad valorem taxes complicates the formula slightly but does not change the main intuition about the role of the elasticities. The elasticities are (typically) defined at the market and good level, which implies that tax incidence should be the same for all firms who compete on the same market.

In models of imperfect competition the role of supply elasticity is replaced by more elaborate assumptions of firm behavior. Instead, the shape of demand curve continues to play an important role. A simplest deviation of perfect competition is to assume a monopoly which sets prices depending on the shape of the demand curve. Weyl and Fabinger (2013) show that the tax incidence for a monopoly can be written in a form that resembles that of perfect competition. In this case, tax incidence is given by

$$\frac{dp}{dt} = \frac{1}{1 + \frac{\epsilon_D - 1}{\epsilon_S} + \frac{1}{\epsilon_{ms}}}$$

where ϵ_{ms} measures curvature of the logarithm of the demand. Thus the shape of the demand largely determines the tax incidence in monopoly case. A zero pass-through would require a perfectly elastic demand. The result also shows that a monopoly would like to operate in a region where demand elasticity exceeds one, thus a zero demand elasticity no longer remain sufficient to get to full pass through. Weyl and Fabinger (2013) further show that in symmetric imperfect competition models the pass-through is something between the perfect competition and monopoly tax incidence. The key lesson is that the demand elasticity and form of imperfect competition determine the price pass-through in the rich class of models previously incorporated in the tax incidence literature. To derive a zero pass-through, one need to assume perfectly elastic demand in these models.

Weyl and Fabinger (2013) also analyze strategic interaction of firms in their context. It is natural to think that competing firm would react to the actions of its competitors, but the possible action space is vast; a firm could change its prices, advertise, differentiate its products and so forth. In general a price decrease in one firm could lead to either price increase or decrease in a competing firm depending on whether firms are strategical complements or substitutes. However, these models do not yield clear predictions for tax incidence, since everything depends on the nature of interaction between firms and is thus left for an empirical analysis to find evidence on this. Despite of this, it is interesting to test in empirical data whether or not firms in the same small geographical areas seem to react to each others price changing decisions due to the reforms.

3 Reforms and data

3.1 The reforms

All counties within the EU use value added taxation (VAT) for consumption taxes. EU regulation stipulate the use of one standard VAT-rate and, at most, two reduced rates.[3] From 2009 an EU Directive (CD 2009/47/EC) allows the countries to apply one of its reduced rates to restaurant services. As a consequence, France, as the first country, reduced its restaurant VAT from 19.6 percent to 5.5 percent in 2009 and some other countries, among them Finland and Sweden, followed a few years later.

In Finland, the VAT-rate for restaurant meals was cut from the standard rate of 22 percent to a reduced rate of 13 percent from July 1st, 2010. In Sweden the corresponding VAT-rate was reduced from 25 to 12 percent starting January 1st, 2012. In both countries the takeaway meals were already at the reduced rate and the reforms did not apply to alcohol, which remained at the standard rate. In both countries the changes in VAT legislation was passed relatively close to the reform, which makes large pre-reform anticipatory effects unlikely.

We measure the the impact on prices by means of the the pass-through, defined as:

$$\Delta = \frac{p^a - p^b}{p^b} * 100 = x\% \tag{1}$$

where p^a is consumer price after the reform and p^b is consumer price before the reform. The full pass-through (i.e. no change in producer prices) would indicate Δ of approximately -7.4% in the Finnish reform and -10.4% in the Swedish reform. Notably, and in contrast to in sales taxes in the US,

[3] Some sectors are zero-rated and some are exempt from VAT.

consumer prices within the EU are always displayed including VAT. Hence, Δ is the price change observed by the consumers on the price tags.

3.2 Outline of the empirical approach

Our basic empirical approach is to compare the price evolution within the restaurants in the countries that were affected by the VAT-reform with the evolution in neighboring countries. We use Estonia as the control for the Finnish reform, and Finland as the control for the Swedish reform. The analysis of the average impact of the reform thus relies on standard differences-in-differences (DD) assumption, i.e. that the behavior of the control group (neighboring countries) should reflect the (counterfactual) evolution of the treatment group in absence of treatment. The rationale for using neighboring countries as controls mimics that of the set of state level DD-studies conducted in a US setting since the minimum wage study of Card and Krueger (1994).[4] As with neighboring states in the US, Finland and Sweden have very similar institutions, geographic location (both share similar climate), share a border, have similar culture, seasonal holidays, vacation periods and seasonality in national food production.[5] They are also covered by the same EU regulations concerning VAT legislation.

Nevertheless, it is possible that the restaurant industries in neighboring countries develop in different ways over time. In our main analysis, we rely on data we collected on our own, starting just before the reform. These data contain a richness (and sample size) that is unavailable in standard CPI-collections of prices, but for obvious reasons they do not cover a very long pre-reform period. To check weather the key assumptions are reasonable, we instead start by illustrating the evolution of the restaurant-meal component of the CPI in Sweden, Finland and Norway (unfortunately we do not have the CPI-data for Estonian restaurant meals). The evolution is shown in Figure 1. As is evident, the CPI meal prices have trends that are largely parallel in the different countries with only two exceptions: Finnish meal

[4]We return to standard concerns regarding standard errors below.

[5]In both countries (as in Estonia) Christmas and New Year are celebrated in similar manner and bank holidays are of similar length and on the same dates.

prices dropped in July 2010 as VAT for Finnish restaurant meals was reduced from 22 to 13 percent, and secondly, Swedish meal prices dropped in January 2012 when VAT was reduced from 25 to 12 percent. Interestingly Finnish restaurant prices seem to catch up the overall trend while Swedish restaurant prices diverge. We provide potential explanations for this difference with our medium-run analysis below.

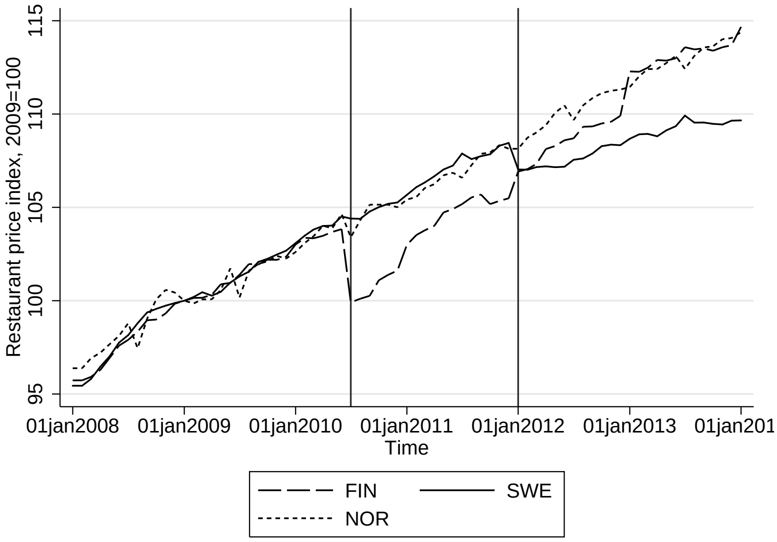

Figure 1: Longer-term development of restaurant prices in Finland, Sweden and Norway

Note: This figure shows the development of monthly consumer price indices collected by statistical offices for restaurant meals in different countries: Finland, Sweden and Norway. The time period is from January 2008 to December 2013 and the index value is set to equal 100 in January 2009. Vertical lines in the figure refer to the VAT cuts for restaurants in Finland (July 2010) and in Sweden (January 2012).

Figure 13 in Appendix provide further evidence along the same lines, but instead focusing on total sales and wage bills within the within Finnish and Swedish restaurant industries. Also these seem to follow each other reasonably well.

3.3 Data

We collected prices directly from the restaurants using our own price collection protocol. We first drew a random sample of restaurants in Sweden, Finland and Estonia from national tax registers. These registers contain all firms liable to taxation in these countries, listed by their primary industry. By using the national tax registers as the base for our random sample, we are also able to to access firm-level register based information on turnover, profits, the number of employees and the total wage bill at an annual frequency from these registers and link these to our survey data.

Data was collected by a separate team of research assistants within each country. Our first choice was to collect the prices from Internet pages of the restaurants. Most, but not all, of the restaurants in the sample did have a website that included prices for meals. If no website was found, we contacted the restaurant by phone. This procedure allowed us to collect prices and other information from a fairly large number of restaurants across a large geographic area based on a random sampling frame.

For each round, we collected prices and meal information for 7 to 11 meals at each restaurant from a defined protocol (depending on the type of restaurant, we collected elements such as starters, main courses, vegetarian meals, pre-set lunch prices and so forth). The assistants chose the exact meals within each category with the intention that these should be possible to follow over time. Since we planned to follow the exact meals across time, it was not essential exactly which meals the assistants chose within each category. Along with the prices, we also recorded other information from restaurants, e.g. restaurant type and other categorical information about the restaurant and the surroundings such as indicators for being located in a mall or on a restaurant-dense street.

In the case of Finnish reform, we collected the pre-reform data in May/June 2010 and the short-run incidence data in July/August 2010. The counterfactual for Finland was chosen to be Estonia (at that stage we were, for obvious reasons, not aware that there would be a reform in Sweden two years later). For the Swedish reform, our pre-reform survey was run in October/November 2011 and the short-run incidence survey in February/March 2012. In the Swedish case, we used Finland as the control country. Although our main analysis focus on the short-run responses, we also repeated the survey half a year and a year and half later, which enables us to also examine medium-term price effects. A disadvantage with the longer term responses is that a larger fraction of meals or restaurants have exited, and thus cannot be followed over time.

3.3.1 Independent restaurants and chains

A main element in our analysis for the tax incidence of the VAT reforms is the role of ownership structure of price-setting firms. Throughout, we define restaurants that (according to our survey) are not part of a chain or franchised restaurants, as *Independent* and other restaurants as *Chains*. More precisely, we define all restaurants belonging to brand names with two or more restaurants as Chains and add restaurants belonging to very large firms (using the top quartile in total firm-level wage bill as the cut off).

We use this split of the data since we conjecture that independent restaurants are less likely to have employees that are specialized on price setting. Naturally, the larger scale of the operation allows the within the Chain-category to be more specialized. In contrast, pricing decisions within independent restaurants are more likely to be made by owners, entrepreneurs or other managers who need to perform a wide set of tasks (including staffing, and possibly, cooking) whereof pricing is just one. The other possible distinction between the two groups could be objective function of the managers. Entrepreneurs in independent firms may want to satisfy their customers, and thus differ from pure profit maximizing behavior. Large chains may have complicated dynamic objectives, such as influencing national tax policies, which would differ from profit maximizing behavior (and that of independent firms) in the short run. Thus, we conjecture that the conditions under which pricing decisions are made are dramatically different for the two types of restaurants.

3.4 Descriptive statistics

Table 1 gives the descriptive statistics divided by the restaurant type. Almost two thirds of the data consist of independent restaurants. The bottom two statistics show that the chain restaurants are much larger in size, are somewhat more often located within a mall and are more likely to belong to an employer confederation. But most other characteristics are surprisingly similar. In particular, the two types contain very similar fractions of fast food restaurants, a la carte restaurants, cafes and lunch restaurants and the average meal prices are only marginally higher in the chains.

In Figure 2 we show the price distributions separately for independent and chain restaurants, divided by treatment status. The figure gives support for the comparability of the treatment and control groups as well as for the two firm types in that the overall shape of price distributions have a substantial overlap. The comparability across treatment status is important for the analysis, since we need to assume that the restaurants in neighboring country represent a counterfactual for the restaurants in the reform country. Although we will rely on differences in differences and therefore do not require exactly the same level of prices before the reform, we take it as positive indication that the price distributions are fairly similar. More importantly in terms of the analysis of restaurant types, however, is the large overlap in price distributions between restaurant types before the reforms. This suggests that the restaurants are competing in roughly similar market segments. In the empirical analysis, we account for remaining differences in pre-reform prices.

3.5 Methods

In the main analysis, we pool across the two reforms, but let the impact vary between independent restaurants and chains as defined above. Our very detailed micro data allow us to follow the development of the price of the same meal served in the same restaurant over time and we are therefore able to analyze $\triangle$ of equation (1) separately by meal. In order to get a uniform measure, we scale the changes relative to the full pass through of each reform. Thus, a change of -1 indicates that the price has changed by

	Chain			Independent		
	Mean	Median	SD	Mean	Median	SD
Share of restaurants	0.369	1		0.631	0	
Price	10.130	8	7.269	8.996	7.304	7.711
Mall	0.192	0	0.394	0.089	0	0.284
Competition	0.160	0	0.367	0.117	0	0.322
Takeaway	0.075	0	0.263	0.066	0	0.248
Confederation	0.573	1	0.495	0.171	0	0.377
Table reservation	0.062	0	0.241	0.034	0	0.181
Changed menu	0.313	0	0.464	0.241	0	0.428
Restaurant classification						
Fast food	0.258	0	0.437	0.223	0	0.417
Ala Carte	0.544	1	0.498	0.553	1	0.497
Cafe	0.072	0	0.258	0.118	0	0.323
Lunch	0.124	0	0.330	0.104	0	0.305
N of firms		791			1,473	
N of prices		4,062			6,954	
Annual wages	22,647,388	1,856,617	75,764,215	315,052	199,581	352,784
Annual turnover	161,839,058	2,361,224	561,524,496	347,890	213,813	450,842

Table 1: Descriptive statistics

Note: The table presents the descriptive statistics of key variables in the data. The summary statistics are divided by restaurant types for Chains and Independents according to our Chain variable. Price is the price of meals in euros. Mall indicates restaurants located in a mall or in a restaurant dense shopping area. Competition refers to restaurants that have another restaurant in the same market-segment very close by. Takeaway indicates restaurants that listed takeaway meals as a possibility in the websites or in the phone survey. Confederation indicates firms belonging to an interest group which aim is to lobby for the hospitality sector (Visita in Sweden and MaRa in Finland). Table reservation equals to one if a customer is able to reserve a table via Internet. Changed menu indicates restaurants that had changed the composition of its menu. Restaurants classifications comprises of for categories: fast food, dinner, cafeterias and pubs in one category and lunch restaurants for those open mainly during lunch and breakfast hours. Annual wages are the total of wage bills and annual turnover is total of tax inclusive sales. The last two variables are come from administrative tax data.

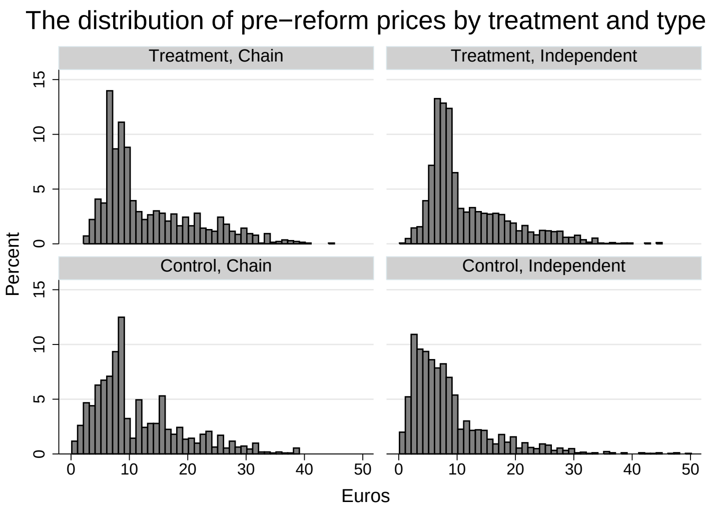

Figure 2: The distribution of pre-reform prices for pooled reforms by treatment and type

Note: The figure shows the distribution of pre-reform meal prices by treatment and type of restaurant. The upper left panel shows this for Chains and the upper right panel for Independents in the treatment group. The lower panels show price distributions divided similarly for the control group.

the amount of the full pass through of the reform in question (-7.4% in the Finnish reform and -11.4% in the Swedish reform). This allows us to analyze the distributional impact in graphical form.

When running DD-regressions for the average impact of the two groups, we let the outcome be the difference in log prices (ΔlnP_{it}) and estimate the following equation:

$$\Delta lnP_{ijr} = \beta_1 D_{jr}^{Treat} + \beta_2 D_{jr}^{Independent} + \beta_3 (D_{jr}^{Independent} * D_{jr}^{Treat}) + \beta_3 (X_{ijr}) + \varepsilon_{ijr}, \tag{2}$$

where the dependent variable ΔlnP is a the difference of logarithmic meal price before and after the reforms of meal i at restaurant j at reform r, where D_{jr}^{Treat} is a dummy for restaurants in the treatment group and $D_{jr}^{Independent}$ is a dummy for independent restaurants. Notably, the difference form takes care of all unobserved meal-specific constant factors. The coefficient β_1 identifies the effect of the VAT reform on change in prices for restaurants that are not independent (i.e. Chains), β_2 measures any additional price trend for Independent restaurants within the control regions and β_3 reveals differences in prices between independents and chains. X contains a vector of other covariates (interacted with treatment status). The role of X is to account for other (market) factors besides ownership structure which could explain differences in tax incidence between the two groups, empirically, we use the variables discussed in Table 1.

To correctly test our hypothesis of whether the estimated pass-through on prices differs from zero, we need to be concerned with the properties of the error term (ε_{ijr}). A standard concern in DD-settings is that the error term may be correlated within groups which needs to be accounted for (see e.g. Bertrand *et al.* 2004). To handle the concern a similar unobserved shock may affecting the errors of all restaurants in one country, we apply block bootstrap method at country level clusters (see Cameron *et al.* 2008 for a discussion).[6]

[6]However, in parts of the analysis, we weight the regressions which prevents us from using the block bootstrap. In these cases, we only robust (heteroscedasticity consistent)

4 Results

In this section we show our empirical results. We start by discussing the overall impact of the reforms on the short-run price change distributions before turning to the analysis where we separate between independent restaurants and chains. We then show some additional evidence on the diverging pricing behavior of independent restaurants and chains. We end the section by discussing the medium term impact of the reforms.

4.1 Overall tax incidence

We first show estimates of the average short-run pass-through of the VAT-reforms onto prices. This impact was already visible in the analysis of the restaurant-meal component of CPI depicted in Figure 1 above. Using our own micro data instead allows us to follow the same meals over time for a large set of data points and to study the anatomy of the price changes. As our final interest lies in the behavior of firms, we do not re-weight our main analysis by sales as is done in the CPI-calculations.

The price change distribution relative to full pass-through is shown in figure 3. The panels depict the treatment and control groups separately (pooled across the two reforms). The large spike at zero indicates that large fraction of meal prices did not change at all in the reform despite of significant reductions in the VAT rate. Naturally, this spike is lower for the treatment group than for control group indicating that the VAT rate reductions had an impact on prices. The second spike in the distribution is around full pass-through for the treatment group, indicating that when meal prices changed, they often changed by the full pass-through. Overall the figure clearly shows the impression of a noticeable, but low overall short-run price effect of the reforms. We return to our estimates of the long-run impact below.

Table 2 quantifies the average short run pass-through of pooled reforms without any controls in column (1) and with additional controls in column (2). Pass-through is the change in prices on average relative to the full pass-

standard errors and indicate how much bias there was in the comparable unweighted regression.

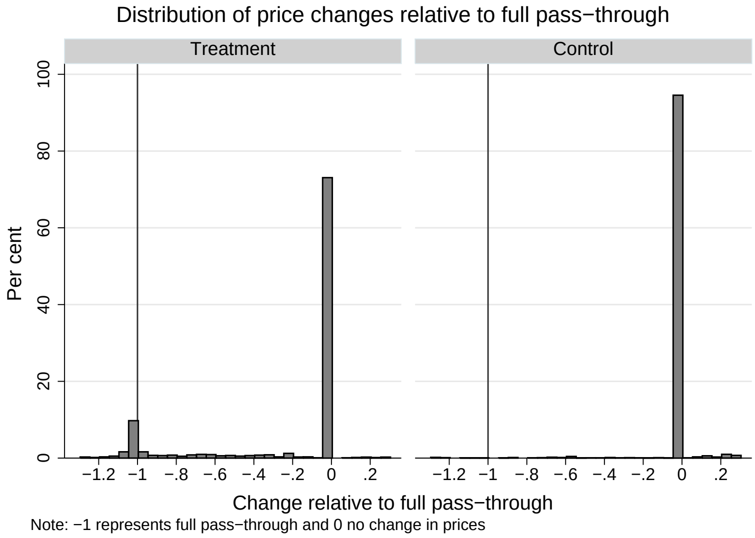

Figure 3: Distribution of price changes in the two reforms
Note: This figure shows the distribution of meal price changes (before vs after) relative to the full pass-through for treatment (left panel) and control restaurants (right panel). Treatment restaurants faced the VAT cuts and control restaurants did not. In the figure, the reforms are pooled together and normalized such that -1 refers to the full pass-through (7.4% price decrease after the Finnish reform, and 10.4% price decrease after the Swedish reform) and 0 refers to no change in prices after the reforms.

through for full pass-through is normalized to be -1 (in percentage terms it was -7.4% decrease in the Finnish reform, and -10.4% decrease in the Swedish reform). Our results imply that the overall pass-through is fairly low, approximately 26 percent of full pass-through in the short run. Different covariates do not seem to have a large impact on the overall pass-through. [7]

4.2 Independent restaurants vs. chains

We now turn to the role of of independent restaurants an chains in explaining the overall price change patterns. Figure 4 shows the price change distributions, separately for chains and independents by treatment status. The pass-through is clearly very different between these restaurant types. About 60%

[7]Table 6 in the Appendix shows the results separately for both reforms.

	(1)	(2)
VARIABLES	Both reforms	Both reforms
Pass-through	-0.271***	-0.259***
	(0.018)	(0.022)
Confederation		-0.106***
		(0.020)
Competition		0.137***
		(0.027)
Changed menu		-0.078***
		(0.018)
Mall		-0.141***
		(0.033)
Constant	0.077***	0.124***
	(0.014)	(0.015)
Observations	6,629	6,629

Block bootstrapped standard errors with restaurant-level clusters and 100 replications

*** p<0.01, ** p<0.05, * p<0.1

Table 2: Average pass-through

Note: The table presents regressions for the change in log prices due to the reforms. The outcome is normalized so that change of -1 would indicate full pass-through. Pass-through is the coefficient of treatment indicator in this regression, indicating the change in prices relative to the full pass-through due to the VAT cuts.

of chain restaurants reduced their prices after the reforms whereas more than 80% of the independent restaurants kept their prices constant despite of the large reductions in VAT rates. Thus, the dichotomy between independents and chains is a key predictor of in which part of the bimodal price change distribution a restaurant ends up.

The patterns for the control group appear more similar as prices do not change at all for most restaurants, which is expected given that short-run nature of the analysis. There is however a (statistically significant) larger probability that chains change their prices in the control state (mostly upwards, for natural reasons). This suggest that independent firms have a less adaptive pricing strategy also in normal times. But, importantly, the low responsiveness of the independents also in absence of the reform does not seem to be the sole determinant of their low average pass through as the mean price change among those that change their prices is substantially lower than for price-changing chain restaurants.

To quantify the short run price changes due to the reforms, we estimate DD-regressions where the outcome is in differences of logarithm form, following equation 2. Table 3 presents the results. The average impact of the reform is around -2.3 percent reduction in meal prices in column (1). In the model with interactions in column (2), where the "treatment" variable captures the impact for independents, we see an impact of -0.009 for the independents. Not surprisingly, the difference to Chains is large (-3.9 percent) and statistically significant.

The purpose of this paper is to see if independent businesses respond differently from chains, for reasons that are related to the their different business models. Hence, Column (3), shows how participation in an employer confederation, a dummy for changed elements in our collected parts of the menu (indicating lower costs of changing prices) and location in a mall affects, all interacted with the reform (main effects are in the model), affects the results. A key contesting hypothesis for the heterogeneity in meal price changes we have seen so far is that they differ in the degree of local competition they face. We use three indicators of the local market: (i) the zip code, (ii) the level of the original (pre-reform) prices, in quartiles, and (iii) restaurant clas-

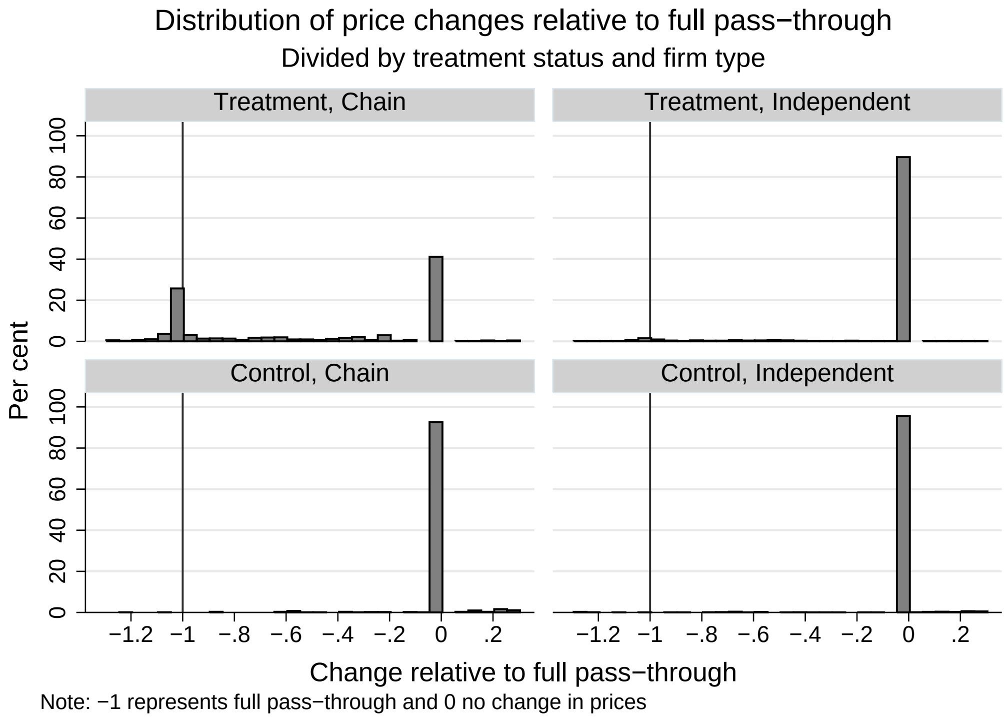

Figure 4: Short-run price changes in divided by treatment status and restaurant type

Note: The figure shows the distribution of meal price changes by restaurant type and treatment status in the reforms relative to the full pass-through. Price changes are normalized such that -1 refers to the full pass-through in each reform and 0 refers to no change in prices. The figure shows the relative price change distribution for all main meals in the data with each price change being one observation.

sification (fast food, a la carte, cafe, lunch restaurant or other). As a first test of the market hypothesis, we have re-estimated the model using zip-code fixed effects (by definition, these vary by treatment status) and (initial) price quartile dummies. As shown in Table 3, column (4) and (5) this has a very marginal impact on the estimate of interest. In column (6) we interact zip code fixed effects with price quartile dummies and restaurant classification indicators. This means that the model only compares across restaurants of different types that compete within the same price range and broad location against similar (broad) types of products. Although the point estimate of the of interest is marginally reduced when adding a very large set of covariates, the main thrust of the difference remain also in these very tight specifications, which we interpret as reassuring.

As a second test, we have looked at the price change distributions, separately by initial price quartile. The results are displayed in Figure 7. As is evident, the difference between independent and chains remain remarkably similar across the distribution. Interestingly, for chain restaurants the graph indicates that the higher is the price quartile the lower is the pass-through.

Our data also contain indicator for being located in a mall. Figure 5 presents results divided according to this variable for the treated restaurants. It is evident that Chain restaurants located in malls do appear to respond more heavily than other chains, which is consistent with our priors, but independent restaurants appear to equally ignore the reform if they are located in a mall or not.

We have also explored the possibility to specifically analyze the restaurants located in specific well-defined restaurant dense locations. So far [this part is to be extended] we only have data on Helsinki, the capital and largest city of Finland, around the Finnish reform. We took areas defined by zip-codes in Helsinki, and divided these areas into smaller areas that define a local competition area. So far we have used these data to analyze how the price change behavior differ depending on whether a neighboring restaurant change their prices or not. The results show high resemblance to the main results presented in figure 4. This highlights that even when located in dense areas and facing competition from neighboring restaurant, Independent restaurants continue to ignore the reform. Intriguingly, even Chain restaurants seem to not to take the behavior of restaurants close by into account.

Overall, we interpret these results as suggesting that neither location, restaurant category or price segments can explain why independent restaurants respond so differently from restaurants belonging to chains. In particular, it seems highly unlikely that similar restaurants located close to each other that serve meals with similar prices before the reforms should face completely different demand elasticities. Furthermore, to explain the zero pass-through for independent restaurants with a conventional model, we should assume that the demand for them is perfectly elastic, which seems even more unlikely.

VARS	(1)	(2)	(3)	(4)	(5)	(6)
Treatment	-0.023***	-0.009***		-0.013***		
	(0.002)	(0.002)		(0.002)		
Chain	-0.013***	0.009***	0.006**	0.011***	0.008**	0.007*
	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.004)
Chain		-0.039***	-0.031***	-0.035***	-0.027***	-0.027***
*Treat		(0.004)	(0.003)	(0.004)	(0.004)	(0.005)
Confederation			-0.016***		-0.014***	-0.019***
*Treat			(0.004)		(0.005)	(0.006)
Changed menu			0.000		-0.001	
*Treat			(0.004)		(0.003)	
Mall			-0.012		-0.011*	-0.003
*Treat			(0.008)		(0.006)	(0.006)
Constant	0.013***	0.005***	0.005*	-0.015*	0.053	0.115***
	(0.001)	(0.001)	(0.003)	(0.009)	(0.038)	(0.027)
N	6,629	6,629	6,629	6,474	6,474	6,474
R^2	0.035			0.126	0.139	0.251
ZIP + PriceQ				Yes	Yes	Yes
ZIP*PriceQ						Yes

Table 3: Variables explaining change in prices

Note: Regressions for change in log prices as an outcome. Column (1) shows has Chain and treatment status as explanatory variables. Column (2) adds to this an interaction between the two variables. Column (3) presents the contribution of some other explanatory variables interacted with treatment status. Column (4) adds zip-code and price quartile fixed effects to the main specification without other covariates and column (5) a similar regression with zip-code and price quartile fixed effects. Column (6) presents a specification where zip-code fixed effects are interacted with price quartile and restaurant classification fixed effects. Standard errors are block bootstrapped at restaurant level.

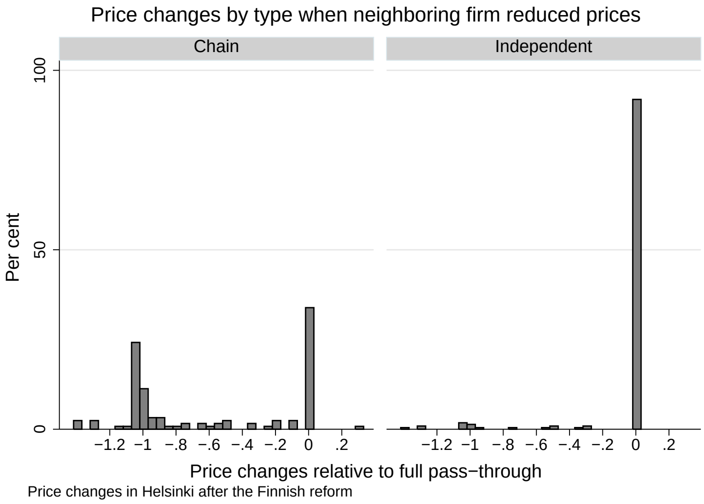

Figure 6: Price changes for those where neighbor reduced prices

Note: The figure shows the distribution of meal price changes relative to the full pass-through for treatment restaurants located in Helsinki whose neighboring restaurant changed at least two meal prices by at least two-thirds of the full pass-through immediately after the Finnish reform divided by our firm type. -1 refers to the full pass-through and 0 refers to no change in prices.

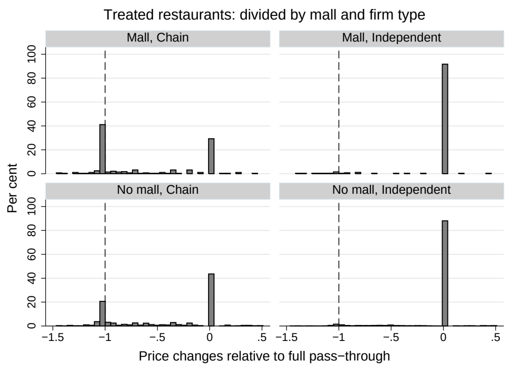

Figure 5: Price changes divided by local competition and located in mall contrasted to ownership structure in treatment group

Note: The figure shows the distribution of meal price changes relative to the full pass-through for treatment restaurants divided by our firm typology and local competition. In horizontal axis -1 refers to the full pass-through and 0 to no change in prices.

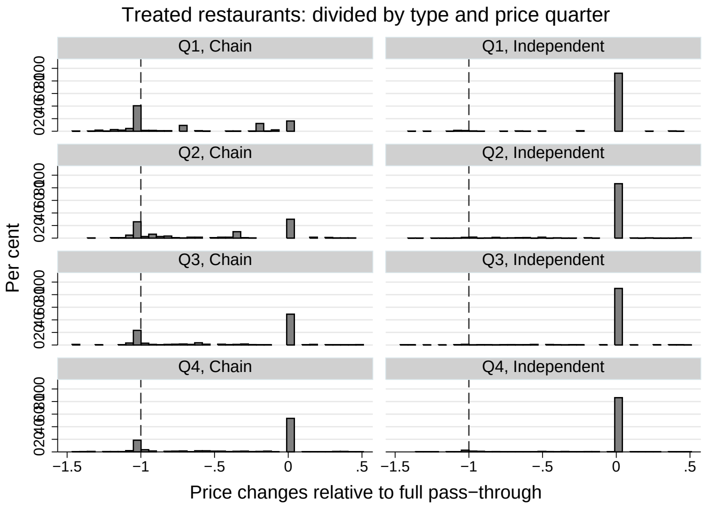

Figure 7: Price changes divided by Price quartile and restaurant type

Note: The figure shows the distribution of meal price changes relative to the full pass-through for treatment restaurants divided by our firm typology and price quartiles. In horizontal axis -1 refers to the full pass-through and 0 to no change in prices

4.3 Other evidence of diverging pricing strategies

4.3.1 Round number pricing

The results presented so far suggest that independent firms are less likely to respond to VAT reforms than restaurants that belong to chains or franchises, even when operating within the same market segment. One possible explanation for this pattern of results is that independent firms have less precise pricing strategies and rely on crude price targets fine-tuned detailed prices. This could be motivated by the complexity involved in figuring out the exact parameters of the demand function that these restaurants are facing (in particular since their economic environment appear to very volatile), or because their objective functions are different. A first piece of suggestive evidence in this direction is provide by the fact that they appear to change prices less often also in normal (non-reform) times as illustrated above.

To provide more evidence on the hypothesis that the independent businesses respond less to the tax reforms because they use cruder pricing rules we have analyzed the the restaurants use of round number prices. The hypothesis being that round number prices is a reflection of less detailed pricing strategies. We define a price as round if it takes an integer value in Euros (in Finland) or 10 SEKs or EEKs (in Sweden and Estonia), which are roughly comparable numbers accounting for exchange rates (all roughly comparable to integer values of USD). Our main interest is in contrasting the incidence of round prices of (e.g. a 9 Euro lunch) to the frequency of close non-round prices (i.e. 8.90 or 9.10 Euro lunches). To this end Figure 8 calculates the distribution of price distances to the closest round number separately for independent restaurants and chains. From the figure it is clear that chain restaurants in the left-hand side panel rely significantly less on round numbers than the independent restaurants depicted in the right-hand side panel. Almost 50 percent of the meal prices are round amongst the chain restaurants whereas the same statistic is less than 20 percent for independent restaurants.

We have also quantified the difference in reliance on round number using regression analysis in order to account for the market segment of the restaurant. Table 4 presents the results from a regression model where the outcome variable is an indicator taking value 1 if the price is round and zero otherwise. This outcome is regressed against multiple covariates, one of which is the Independent dummy. Although several of the covariates explain the round number pricing, the most significant and largest point estimate we find is for the indicator for independent restaurants.[8] Independent restaurants are 12 percentage points more likely to use round number prices than chain restaurants and the differences remain stable and statistically significant when more covariates are added.

[8]Restaurants with local competition have less round prices, and the same applies for restaurants belonging to the employer confederation and those to those who changed some content of their menu. Restaurants located in malls and fast food restaurants appear to use more round meal prices.

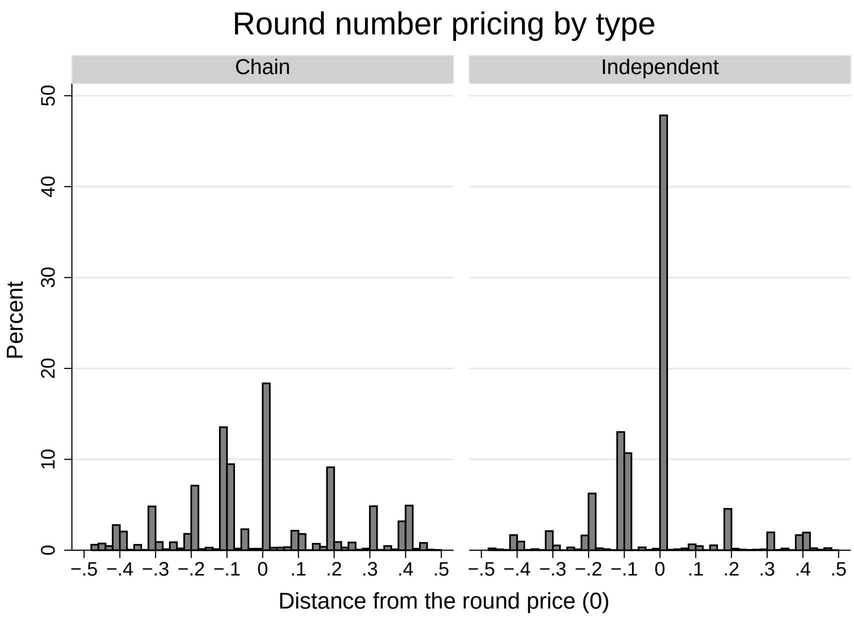

Figure 8: Round number pricing by type

Note: The figure shows the distribution of price distances to the closest round number by our restaurant typology. The figure defines a restaurant meal price to be a round number among Finnish restaurants if it is an integer value in euro, for Swedish restaurants if it is multiplicative of 10 SEKs and for Estonian restaurants if it is multiplicative of 10 EEKs. (The exchange rates of 1 euro = 9.06 SEK = 15.65 EEK in December 2010). In the figure, the round price is normalized to zero and the bandwidth is 0.05 units.

VARS	Round	Round	Round
Chain	-0.115***	-0.121***	-0.126***
	(0.007)	(0.007)	(0.007)
Right after	-0.006***	-0.006**	-0.006***
	(0.002)	(0.002)	(0.002)
3-6 months after	-0.047***	-0.046***	-0.056***
	(0.005)	(0.005)	(0.005)
15-18 months after	-0.012**	-0.010*	-0.020***
	(0.006)	(0.005)	(0.005)
Confederation	-0.078***	-0.068***	-0.045***
	(0.007)	(0.006)	(0.008)
Competition	-0.053***	-0.066***	-0.050***
	(0.009)	(0.009)	(0.009)
Changing menu			-0.042***
			(0.007)
Mall			0.056***
			(0.010)
Country (ref. Finland)			
Sweden			-0.016*
			(0.009)
Estonia			0.067***
			(0.011)
Constant	0.283***	0.289***	0.297***
	(0.006)	(0.005)	(0.009)
Observations	34,874	34,874	34,874

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 4: Round number pricing:

Note: Regressions for an indicator of round number price as an outcome. The main variable of interest is the Chain variable introduced in column (1). Column (2) and (3) introduce more covariates shown in table.

4.3.2 Strategic price increases during currency conversions

The results presented above showed that a large share of the (mostly chain) restaurants that actually responded to the reforms, did so by allowing for a full pass-through onto prices. This behavior is difficult to reconcile with standard tax incidence theories, but it could be viewed as an outcome of strategic price setting behavior if the firms believed that their pricing behavior was monitored either by customers, who may react particularly positive to price reductions that hit the full pass-through (anecdotes from advertisements suggest that this may have been the case), or because large firms coordinated their short-run price response in order to prevent the governments to reset the VAT-rates.

In order to find external evidence on how strategic price changing behavior may differ between independents and chains, we have analyzed the price responses to the conversion from Estonian Krooni (EEK) to Euro. This is an interesting experiments since changing the name of the currency is expected to leave marginal production costs unchanged, and only require a change of price tags. On the other hand, customers may find it difficult to keep track of the exact prices during the conversion. Thus, it potentially creates an opportunity for firms to strategically increase their prices without negative customer reactions. Our conjecture is that chains should use this opportunity more that independent restaurants if the chains, as we believe, are more strategic in their price setting behavior.

The resulting relative price change distributions are shown in figure 9. Each panel shows the relative price changes across two collection moments at different time intervals. The results show that restaurants belonging to chains (relative to independents) increased their prices more often right at the time of the currency conversion than in surrounding time periods. According to standard theories changing the currency should not affect prices, and taking that as an opportunity to increase prices could well benefit the restaurants which do so. Thus, we interpret the evidence as consistent with the notion that chains rely more on strategic pricing behavior than independent restaurants.

4.3.3 Business volatility

As a final piece of evidence supporting our division into two types of restaurants, we show how volatile the market environment is for restaurants and

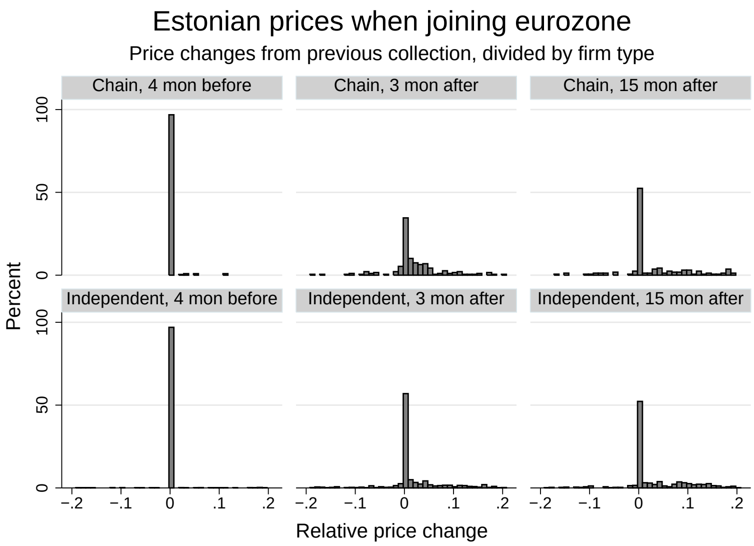

Figure 9: Price changes around Estonian currency change in the two restaurant groups

Note: This figure shows the distribution of longer term meal price changes for Estonian restaurants by our restaurant typology (chain in the upper panel and independent in the lower panel) before and after Estonia joined the Eurozone. The first column of panels show price changes from 7 to 4 months before joining the Eurozone, the middle column shows changes from 4 months before to 3 months after and the last column from 3months after to 15 months after joining the Eurozone. The horizontal axis shows relative price change in percentage terms.

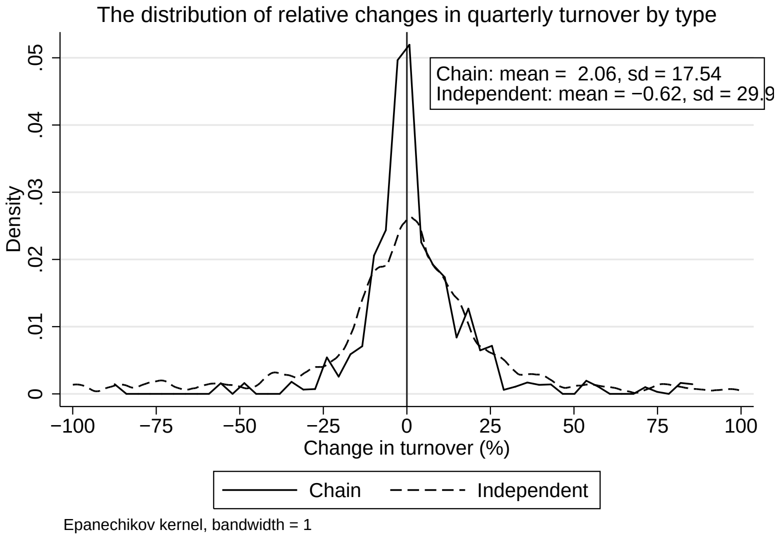

Figure 10: Changes in sales relative to own history

Note: This figure shows the distribution of kernel densities of relative changes in quarterly sales for chain and independent restaurants. We calculate for each firm a relative change in sales from two quarters before and after the reforms. In the figure, we restrict the changes to be between -100 and 100 percent. The bandwidth is 1 percent.

also how it differs by the type of restaurants. In figure 10 we illustrate the kernel density of the distribution of changes in quarterly total sales relative to sales that the same restaurant had in the same quarter last year. Thus this method controls for firm-level heterogeneity and seasonality within a year. Despite of these controls, there remains huge variation in the distribution of changes in total sales. This indicates that restaurants need to operate in a market environment where the volume of purchases and number of employees need to be constantly adjusted. More importantly for the analysis here, the distribution of changes in turnover is wider for independents than for chain restaurants. This means that the independent entrepreneur or manager who may already be burdened by many strategic choices, needs to pay attention to larger firm level shocks than the price setting specialist employed by chain restaurants.

4.4 Medium run tax incidence

So far, we have focused entirely on the short-run effects on prices. We now turn to the longer run effects using data from four separate collections: the first (as before) 1-2 months before the reforms, the second (again, as before) 1-2 months after the reforms, the third is 3-6 months after the reforms and the fourth is 15-18 months after the reforms. We follow the change in the same meal price over time. Obviously, some of the meals have changed, reducing the sample size as the time from the first collection elapses. On the other hand, following the same meals allows us to provide specific measures on the changes in prices and control for the unobserved meal size and quality.

We have quantified the average price changes over time in regressions based on equation 2. Table 5 displays the regression results. The results indicate that the immediate reduction in prices is about . The first column shows that the difference between the types in the treatment group starts to diminish over time because chain restaurants first decreased, and later increased, their meal prices relative to independent restaurants.

Figure 11 shows the distribution of meal price changes between the first collection and the consecutive three collections for the treatment group. The upper panel of the figure is for chains and the lower panel for independent restaurants. The first panels from the left is the immediate price change (the same as in figure 4) , the second panels the the prices 3-6 months after the reform and the third panels 15 to 18 months after the reform. The figure shows that the initial full pass-through in chain restaurants vanishes almost completely already within 6 months from the reform. The figure also indicates that a non-trivial fraction of meal prices remained stable during a one and a half year observation period in both types of restaurants.

As a final exercise, it turns out to be illustrative to separate the longer run price responses depending on whether the initial price was changed or not, despite the obvious endogeneity. Figure 12 shows the results from this exercise. It turns out that many of the prices that straight after the reform were at full pass-through were reverted over time. Moreover, the spike at zero starts to reappear (i..e a non-trivial fraction of prices were changed

	Without control group	With control group
Right after	-0.030***	-0.039***
* Chain	(0.002)	(0.003)
3-6 months after	-0.022***	-0.035***
* Chain	(0.002)	(0.005)
15-18 months after	-0.012**	-0.032***
* Chain	(0.005)	(0.009)
Right after	-0.005***	0.004***
	(0.001)	(0.001)
3-6 months after	0.000	0.008***
	(0.001)	(0.003)
15-18 months after	0.029***	0.033***
	(0.003)	(0.005)
N	13,892	22,513
R^2	0.069	0.058

Table 5: Longer run price responses by type

Note: The table show regression result for changes in log prices relative to the pre-reform prices as outcome. The table shows the development of prices in different price collections, 2 months after, 3-6 months after and 15-18 months after. These are shown independently and interacted with the Chain variable.

back to their pre-reform levels). Although the independent firms look more similar to the chains in this dimension it is important to note that there very few independent restaurants allowed for the full pass-through initially. As a contrast, the bottom figure shows the price change patterns to those meals which price initially was not reduced and these prices remained much more stable - and when price changes occurred, they were mostly in the form of increases.

5 Conclusions

The previous literature on tax incidence has been heavily focused on situations where the tax incidence only depends on market level conditions such as the elasticities of demand and supply, and the degree of competition

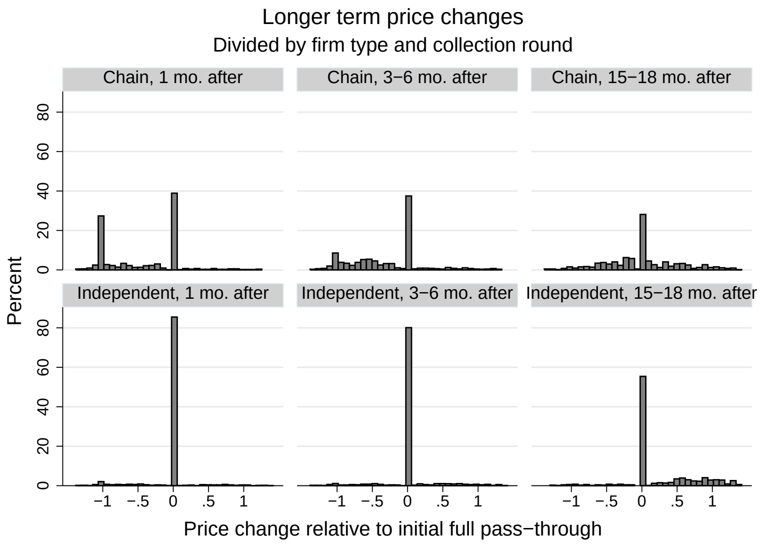

Figure 11: Longer-run price changes by firm type

Note: The figure shows the distribution of longer term meal price changes relative to the initial full pass-through solely for treatment restaurants by our restaurant typology. In the horizontal axis -1 refers to the initial full pass-through and 0 refers to no change in prices right after the reforms. The figure offers price changes 1 month after the reforms (left panels), 3-6 months after the reforms (middle panels) and 15-18 months after the reforms (right panels) as a change from the pre-reform prices.

(e.g.Weyl and Fabinger 2013).[9] In this paper we have instead documented that different types of firms respond very differently to consumption tax reforms.

Our results from two separate substantial VAT reductions in Sweden and Finland show that the overall immediate pass-through pattern was bi-modal. Many meal prices remained constant in the short-run and others were reduced by the exact amount corresponding to a full pass-through. Restaurant ownership structure explains a significant part of this pattern. Almost all of the independent restaurants kept their prices constant and thus effectively ignored the reform. Notably, standard models including both perfect competition case and and many imperfect competition scenarios would only predict

[9] Kopczuk et al. (2013) enrich this view by taking into account the fact that firms may have different opportunities to evade taxes depending on, for example, where they are located in the supply chain.

a zero pass-through if demand is perfectly inelastic (Weyl and Fabinger 2013).

Contrary to the finding for independent group, a substantial fraction of restaurants belonging to chains or franchises reduced their prices to a full pass-through during the reforms. Accounting for very detailed indicators of market segments such as price location and restaurant category does not explain the difference between restaurant types.

Given that independent restaurants are likely to be run by entrepreneurs or managers who need to concentrate on many other tasks than just pricing strategies, one possible explanation is that these firms have limited attention to pricing decisions. Thus, they could behave in a way which is more akin to non-professionals (i.e. consumers) who have been found to exhibit limited attention to taxes in previous studies. if the managers of independent restaurants are constrained in the amount of time, talent or interest they allocate to pricing decisions they may resort to less-complicated pricing strategies. Several of our empirical facts support the notion of constrained optimization. Apart from the low impact of the reform, we also find that they are less likely to change their prices, rely more on round number pricing and were less likely to use the opportunity to raise prices that were given by the Estonian currency conversion when adapting the Euro.

It is notable that the pass-through pattern for restaurants belonging to chains on average is more in line with expectations from standard models in the sense that they reduced their prices in the response to the reform. What is less expected is that, within this restaurant class, some restaurants shifted the tax reduction to prices fully, while others did not react at all to the reforms. This leads to an expected average pass-through, but the dichotomous anatomy of these price changes does not follow directly from standard theory. One possible explanation is that restaurants that responded by a full pass-through did so for strategic reasons, either relative to customer responses or as part of a coordinated effort to ensure that policy makers kept the reduced VAT rates in the future. The fact that that many of the restaurants with a full short-run pass-through increased them only 3 to 6 month after is consistent with this explanation.

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Appendix

VARIABLES	(1) FIN reform	(2) FIN reform	(3) SWE reform	(4) SWE reform
Pass-through	-0.311***	-0.182***	-0.175***	-0.180***
	(0.025)	(0.030)	(0.023)	(0.025)
Confederation		-0.326***		-0.022
		(0.030)		(0.027)
Competition		0.054		0.060*
		(0.037)		(0.031)
Changed menu		-0.067***		-0.024
		(0.024)		(0.029)
Mall		-0.084**		-0.118**
		(0.043)		(0.047)
Constant	0.002	0.024	0.135***	0.150***
	(0.020)	(0.018)	(0.017)	(0.022)
Observations	3,376	3,376	3,253	3,253

Block bootstrapped standard errors with restaurant-level clusters and 100 replications

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 6: Average pass-through

Note: The table presents regressions for the change in log prices by reforms, columns 1 and 2 are for the Finnish reform and columns 3 and 4 are for the Swedish reform. The outcome is normalized so that change of -1 would indicate full pass-through. Pass-through is the coefficient of treatment indicator in this regression, indicating the change in prices relative to the full pass-through due to the VAT cuts. Columns 1 and 3 offer results from specifications without any control variables, and columns 2 and 4 adds control variables that are visible from the table.

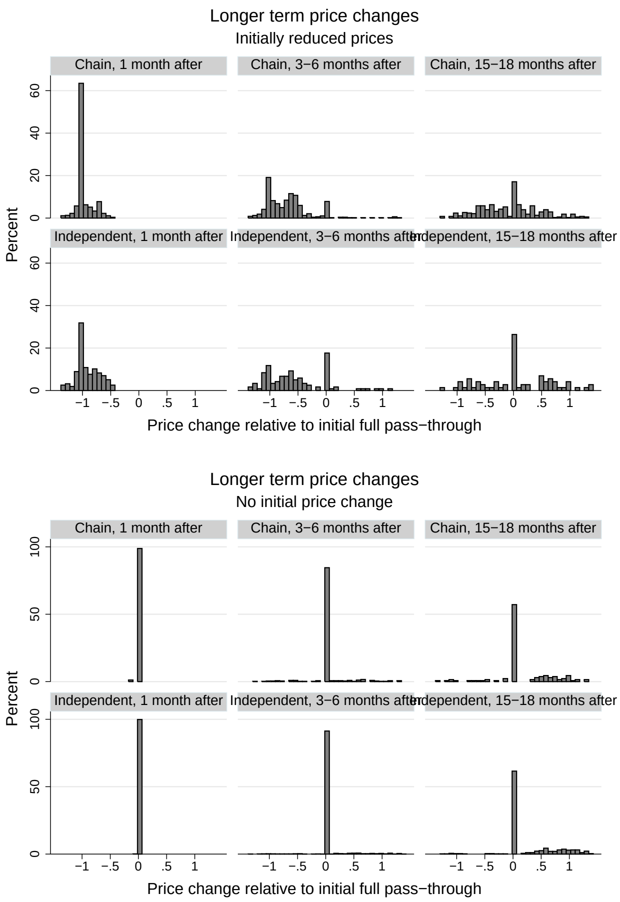

Figure 12: Longer term price changes according to whether initially reduced prices or no

Note: The figure shows the distribution of longer term meal price changes relative to the initial full pass-through solely for treatment restaurants by our restaurant typology. The upper figure is for those that initially reduced prices and the lower figure for those that initially did not reduce prices. In the horizontal axis -1 refers to the initial full pass-through and 0 refers to no change in prices right after the reforms. The figure offers price changes 1 month after the reforms (left panels), 3-6 months after the reforms (middle panels) and 15-18 months after the reforms (right panels) as a change from the pre-reform prices.

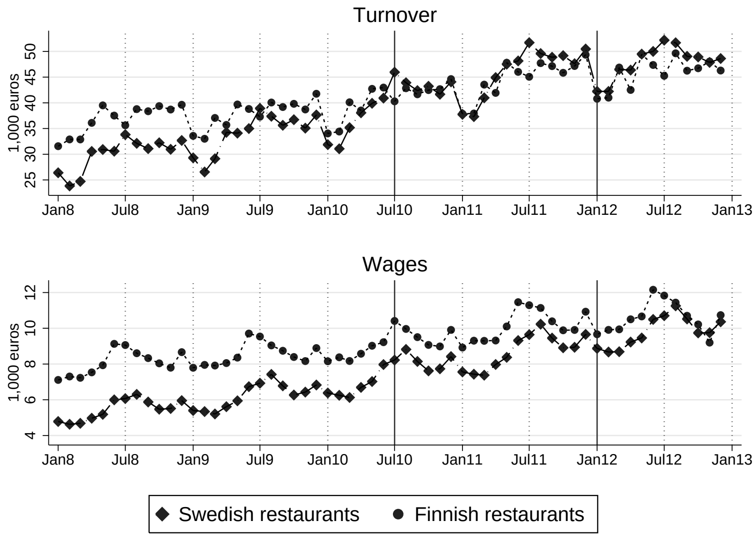

Figure 13: Longer-term development of tax inclusive turnover of restaurants in Finland and Sweden

Note: The figure shows the development of monthly tax inclusive turnover (sales) in the upper panel and wage sums paid to employees in the lower panel of Finnish and Swedish restaurants from January 2008 to December 2012 in thousands of euros. Vertical lines in the figure refer to the VAT cuts for restaurants in Finland (July 2010) and in Sweden (January 2012).