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Passthrough of Retail Price Regulation in the Market for Pharmaceuticals

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Abstract

I study the passthrough of regulated ad valorem pharmacy markups and reduced Value Added Tax (VAT) rates to pharmaceutical retail prices in Finland. My reduced form evidence shows that pharmaceutical manufacturers respond to a decrease in regulated pharmacy markups by increasing their wholesale prices. I estimate a structural model of pharmaceutical supply and demand using data from the Finnish statin market. My results show that manufacturers benefit significantly from existing pharmaceutical tax subsidies, making them an expensive way to support poorer patients.

Tiivistelmä

Lääketaksasääntelyn siirtyminen hintoihin lääkemarkkinoilla

Tässä tutkimuksessa tarkastellaan lääketaksan ja arvonlisäverotuksen siirtymistä lääkkeiden hintoihin Suomessa. Tulosten perusteella lääketaksan alentaminen nostaa lääkkeiden tukkumyyntihintoja. Tutkimuksessa mallinnetaan myös lääkkeiden kysyntää ja tarjontaa kolesterolilääkemarkkinoilla. Tulosten mukaan lääkeyhtiöt hyötyvät merkittävästi lääkkeiden alennetusta arvonlisäverokannasta, mikä tekee verotuesta tehottoman tavan tukea lääkkeiden kysyntää.

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Keywords: Pharmaceuticals, Passthrough, Price regulation, Commodity taxation, Market power

Asiasanat: Lääkkeet, Hintasääntely, Arvonlisäverotus, Markkinavoima

JEL: I11, H22, H25, H51, L51, C23

Reduced Value Added Tax (VAT) rates on pharmaceuticals are widely employed across Europe with the intention of improving consumer access to essential medicines by making them more affordable. Currently, 24 out of 27 European Union (EU) member states utilize reduced VAT rates for pharmaceuticals, highlighting the common belief that such tax subsidies are fully passed through to consumers. However, an important and less discussed question arises: What if these well-intentioned VAT subsidies inadvertently increase costs by enabling pharmaceutical manufacturers to capture a significant portion of these tax benefits? This concern is especially pressing given that Europe's aging demographics drive up demand and strain already constrained public insurance systems.

Similar unintended consequences can arise from the regulation of retail pharmacy markups, a practice prevalent in countries such as Finland, Sweden, Spain, Belgium, and Germany. However, from the perspective of pharmaceutical manufacturers, government-imposed retail markups are no different from a VAT because both are calculated as ad valorem rates—percentages added on top of the wholesale price—that raise the final consumer price without increasing the revenue received by the manufacturer. Unlike VAT, these regulated retail markups directly transfer revenues to privately owned pharmacies rather than contributing to public funds. This paper investigates the extent to which reduced VAT brackets and regulated pharmacy markups can paradoxically lead to higher overall costs from a societal perspective.

I study the passthrough of regulated pharmacy markups and VAT rates to pharmaceutical retail prices in Finland. Accurate estimates of the passthrough of taxes or regulation to consumer prices are a crucial aspect of policy evaluation and market design. For example, reduced VAT rates are often intended to subsidize demand, but incomplete passthrough may instead result in unintended subsidies to suppliers (Kosonen 2015). Despite this, policymakers often do not explicitly consider passthrough, or they operate under the assumption of complete passthrough.[1] A common naive approach is to account for behavioral responses on the demand side using elasticities, but this would still neglect producer responses.

I use two empirical frameworks to study passthrough. I start with a reduced-form model where I estimate the passthrough of a decrease in pharmacy markups to retail prices, utilizing the fact that, due to regulation, not all producers were able to increase their wholesale prices to offset the decrease in pharmacy markups. The reform in question occurred in Finland in 2014. The results suggest that the passthrough was on average only 28%—implying that pharmaceutical manufacturers were able to capture most of the decrease in retail markups by increasing their own wholesale prices. Equivalently, this implies an incomplete passthrough of ad valorem taxes to retail prices.

1. For example, the last two government proposals on markup regulation in Finland do not discuss or mention the effects of pharmacy markup on wholesale prices or vertical market structure. See HE 245/2022 vp and HE 170/2013 for the government proposals (in Finnish).

I also estimate a structural model of supply and demand using data from the Finnish cholesterol drug market, which allows me to analyze counterfactual tax reforms that have not yet been implemented. The demand model is a random utility discrete choice model, and the supply model is based on Bertrand-Nash competition with retail price regulation. In line with my reduced-form results, I find incomplete passthrough for ad valorem markup regulation and taxes. My estimates yield an average passthrough rate of approximately 58%.[2] Consistent with results from the existing literature, I find that firms with higher markups have smaller passthrough rates (Miravete, Seim, and Thurk 2018; Pless and Benthem 2019; Genakos and Pagliero 2022). My structural estimates imply that manufacturers benefited significantly from the policy change, increasing their profits by almost 6% during the years 2014–2017 in the statin market alone. Consumers and the public sector experienced reductions in pharmaceutical expenditure. Since pharmacy profits decreased, the decrease in pharmacy markups was a transfer of rents from downstream pharmacies to consumers and to upstream drug manufacturers.

Motivated by EU regulations that allow Finland VAT rates of 10%, 14% or 24%, I simulate two additional scenarios with higher rates VAT alongside the 2014 markup reform. Compared to the original reform, manufacturers and pharmacies incur losses, but the government gains substantially, despite increased consumer spending.

My results are especially important for small open economies without a significant domestic pharmaceutical industry and with a generous public insurance system, such as Finland and other similar European countries. In these countries, most pharmaceuticals are sourced from abroad. Due to inelastic demand induced by the reimbursement system, wholesale prices are a good proxy for the social cost of pharmaceutical care. Reduced VAT rates are a costly method of reducing consumer expenditures because a significant part of the tax incidence falls on the supply side. This means higher wholesale prices than without the tax subsidy. Removing these would cause smaller price responses than predicted by the canonical tax incidence model. Importantly, reimbursement systems already mitigate the negative effects on consumer welfare by insuring the most vulnerable consumers against price increases.

More generally, the regulator should pay attention to the vertical market structure between drug manufacturers and the pharmacy retail sector, because the retail price regulation is passed through to wholesale prices. Cost control policies in the retail sector, such as markup regulation, can increase wholesale prices. I show that the policy maker can take this into account by preferring VAT as a policy tool to lower wholesale prices, especially if the government reimburses most of the pharmaceutical costs.

This paper is related to three different strands of literature. First, it is related to

2. Because my reduced-form estimation sample consists of different products and different markets, these two passthrough estimates are not exactly comparable.

the literature on passthrough and tax incidence,[3] with an application to pharmaceutical markets and regulation. A key theoretical result, demonstrated by Weyl and Fabinger (2013), show that under imperfect competition, the level of passthrough depends not only on the demand elasticities—like in models of perfect competition in earlier papers, summarized in e.g. Myles (1995)—but also on the curvature of the demand. In general, there has been a growing interest in studying firm responses to taxation. For example, Benzarti et al. (2020) study the assymetry of tax responses in the supply side, showing that firms respond more to tax increases than to tax cuts. Most importantly for my work, Miravete, Seim, and Thurk (2018) show that ad valorem taxes are often strategic substitutes for wholesale prices in differentiated product markets with oligopolistic competition. This implies that changes in taxes (or regulated retail markups) are offset by firms' behavioral responses, which shift tax incidence away from the demand side.

I contribute to the literature by showing how pharmaceutical price controls can act as a form of taxation through regulation. However, compared to price regulation, VAT has several advantages. First, through incomplete passthrough, it lowers wholesale prices, and second, it generates government tax revenue. While it also raises retail prices, this negative effect can, in theory, be offset by adjusting transfers within the social insurance system. My results question the common practice of applying reduced tax rates to pharmaceuticals to reduce consumer prices and expenditure. VAT can serve as a mechanism to offset the effects of market power and imperfect competition by reducing wholesale prices.

I also contribute to the reduced-form literature studying the effects of regulation in pharmaceutical markets. Most existing research has focused on studying the effects of consumer choice policies and regulation on pharmaceutical prices and expenditure (Pavcnik 2002; Brekke, Grasdal, and Holmås 2009; Brekke, Holmas, and Straume 2011; Kortelainen et al. 2023). However, Danzon and Chao (2000) presents some descriptive evidence that the regulation of pharmacy markups undermines competition and the savings potential of generic competition. The results from both my reduced-form and structural models verify that the regulation of pharmacy markups also affects pharmaceutical manufacturers and competition.

This paper is also related to the structural estimation of pharmaceutical demand (Duso, Herr, and Suppliet 2014; Kaiser et al. 2014; Dubois and Sæthre 2020; Dubois, Gandhi, and Vasserman 2022; Atal, Cuesta, and Sæthre 2022; Janssen 2023). I contribute to the literature by modeling the vertical market structure under strict retail price controls established by the government. Dubois and Sæthre (2020) study the effects of parallel trade on negotiations between pharmaceutical manufacturers and pharmacy chains in

3. Notable examples from this literature include Wang (2015) (soda taxes) Duggan, Starc, and Vabson (2016) (health insurance) Hong and Li (2017) (grocery retail) Conlon and Rao (2020) (excise taxes) Kosonen (2015), Harju, Kosonen, and Skans (2018), and Benzarti et al. (2020) (VAT) Hollenbeck and Uetake (2021) (marijuana taxes) Genakos and Pagliero (2022) and Harju et al. (2022) (gasoline) Galloway and Li (2023) (food subsidies).

Table 1: Retail prices for RX drugs in Finland

Wholesale price (WP)	Retail price (2003)	Retail price (2014)	Retail price (2023)
0–9.25 / 0–7.49	1.5 × WP + 0.50 €	1.45 × WP	1.42 × WP
9.26–46.25 / 7.50–39.99	1.4 × WP + 1.43 €	1.35 × WP + 0.92 €	1.35 × WP + 0.52 €
46.26–100.91 / 40.00–99.99	1.3 × WP + 6.05 €	1.25 × WP + 5.54 €	1.24 × WP + 4.92 €
100.92–420.47 / 100.00–399.99	1.2 × WP + 16.15 €	1.15 × WP + 15.63 €	1.15 × WP + 13.92 €
over 420.47 / 400.00–1499.99	1.125 × WP + 47.68 €	1.1 × WP + 36.65 €	1.10 × WP + 33.92 €
over 1 500			1 × WP + 183.92 €

Notes: This table presents the markup regulation for RX and OTC pharmaceuticals in Finland. The first column gives the brackets used in 2003–2022 on the left and the brackets for 2023 and onwards on the right. The second column the retail price formulas applied to RX products between 2003–2013 and for OTC products between 2003–April 2022, after which they apply as maximum pharmacy markups. The third column gives the RX formulas for 2014–2022 and the fourth column presents the current markup formula for RX drugs. The retail prices here do not include the VAT.

Norway. In my application, there is no bargaining between the upstream and downstream firms because of government regulations; the manufacturer is tied to uniform prices throughout the country. This significantly simplifies the estimation of tax passthrough.

The remainder of this paper is structured as follows. I summarize the regulatory environment of the Finnish pharmaceutical market in Section I. Section II gives an overview of my data. I present the descriptive reduced-form evidence in Section III. Section IV introduces my structural model for the statin market, and Section V presents the results. I offer my conclusions in Section VI.

I Institutional Background

A The Finnish market for pharmaceuticals

I study the Finnish pharmaceutical market, where an extensive public insurance system covers more than two-thirds of total expenditure. The market can be characterized by a vertical supply chain where upstream manufacturers set their wholesale prices at the national level (uniform pricing) and downstream retailers (pharmacies or pharmacists) distribute the drugs to consumers. Pharmacies do not set their own prices, but instead the government regulates the retail prices of all pharmaceuticals as a piecewise linear function of wholesale prices. The government also collects 10% VAT on the retail price.[4] Table 1 presents pharmacy markups between 2003–2013 and 2014–2023 for prescription (RX) drugs. Markups for RX drugs were cut in 2014. [5]. The markups of over-the-counter (OTC) drugs are also regulated by the government and were in the 2003 RX regime (Table 1, Column 2) until 2022 when the markup rule changed from a binding formula to the maximum markup, thus allowing price competition for OTC drugs.

4. The tax rate for pharmaceuticals was 9% before 2012, see Value Added Tax Act 1202/2011.
5. The cuts were slightly larger in relative magnitude for more expensive pharmaceuticals

Finland has a public reimbursement system for pharmaceuticals. The reimbursement rate varies from 40% to 100%, depending on the severity of the disease for which the drug is used. There also exists an annual expenditure cap, after which the consumer is fully reimbursed except for a small fixed co-payment per prescription. Wholesale prices of publicly reimbursed pharmaceuticals are subject to price caps that are negotiated between the government and pharmaceutical manufacturers.[6] These price caps are the main regulatory tool to control the costs of publicly reimbursed on patent drugs. However, it should be noted that these price caps are only part of the reimbursement system. Any company that has market authorization from the Finnish Medicines Agency (Fimea) or the European Medicines Agency (EMA), can sell their product at any price they like if it is not included in the reimbursement system.

At the time of loss of exclusivity and the start of generic competition, the former price caps are no longer renegotiated, but they remain in place. However, these products are then subject to reference pricing, under which the government caps the level of reimbursement to a reference price, which is based on the lowest price within a set of substitutable products. The substitution groups are determined by Fimea. These reference prices are set and updated quarterly.

In my reduced-form estimations, I take advantage of the fact that some products were subject to binding price caps at the time of the markup reform 2014. For these products, pharmaceutical manufacturers were unable to increase their wholesale prices to benefit from the reform. On the other hand, for all those products that had no price caps or whose price caps were not binding, firms were able to partly or even fully capture the change in retail prices induced by the reform. In theory, companies could set their new wholesale price at the exact level where retail prices remained constant. Thus, the existence of these two groups of products—separated by the price cap regulation—creates a quasi-experimental setting to study the passthrough of the markup regulation to wholesale prices.

B Pharmacy markup regulation in the European Union

In this subsection, I briefly discuss the regulatory environment in the EU single market with regard to pharmacy markup regulation and VAT rates. This discussion is motivated by the fact that many EU countries mandate pharmacy markups and offer subsidized VAT rates for pharmaceuticals. The overview presented in this subsection highlights the broader relevance of my results beyond Finland. Table 2 provides an overview of pharmacy markup regulation in EU countries, listed in descending order of pharmaceutical market value, with Germany being the largest market.

6. Notice that due to regulation wholesale prices are actual transaction prices and not estimates, so there should be no concerns over measurement errors.

Table 2: Pharmacy Markup Regulations in the EU

Country	PRP formula	PRP cap	Market value	VAT (RX)	VAT (OTC)	Standard VAT
Germany	Yes	Yes	42 962	19%	19%	19%
France	No	Yes	29 552	2.1%	10%	20%
Italy	No	Yes	23 446	10%	10%	22%
Spain	Yes	Yes	17 604	4%	4%	21%
Poland	Yes	Yes	7 239	8%	8%	23%
Belgium	Yes	Yes	6 303	6%	6%	21%
Netherlands	No	Yes	6 185	9%	9%	21%
Greece	Yes	Yes	5 381	6%	6–13%	24%
Austria	No	Yes	4 827	10%	10%	20%
Sweden	Yes	Yes	4 570	0%	25%	25%
Romania	No	Yes	4 500	9%	19%	19%
Portugal	No	Yes	3 524	6%	6%	19%
Czech Republic	No	Yes	3 389	10%	10%	21%
Denmark	Yes	Yes	3 243	25%	25%	25%
Finland	Yes	Yes	2 762	10%	10%	24%
Hungary	Yes	Yes	2 558	5%	5%	27%
Ireland	No	No	2 354	0–23%	0–23%	23%
Slovakia	Yes	Yes	1 461	10%	20%	20%
Bulgaria	No	Yes	1 414	20%	20%	20%
Croatia	No	Yes	1 036	5%	5%	25%
Lithuania	No	Yes	866	5%	21%	21%
Slovenia	No	Yes	743	9.5%	9.5%	22%
Estonia	No	Yes	359	9%	9%	20%
Latvia	No	Yes	275	12%	12%	21%
Malta	No	No	196	0%	0%	18%
Luxembourg	Yes	Yes	184	3%	3%	17%
Cyprus	No	No	177	5%	5%	19%
% ´Yes´	41%	89%	-	-	-	- %
Total	11	24	177110	-	-	- %

Notes: The first two columns indicate whether a country uses a formula to determine the retail price in pharmacies and whether there is a cap on pharmacy margins or prices. The third column displays the pharmaceutical market value in millions of euros for the year 2020. The last three columns show VAT rates for RX and OTC pharmaceuticals and the standard VAT rate. In Ireland, the VAT is 0% for oral medications and 23% for others.

Column 2 shows which EU countries use a pricing formula to determine Pharmacy Retail Prices (PRPs), as is the case in Finland, and Column 3 shows which countries have a price cap on PRPs. The column 'PRP formula' shows if government regulations directly set the retail markups for pharmaceuticals, resulting in uniform prices across all pharmacies. Most countries that do not directly regulate pharmacy prices regulate them with price caps; Ireland, Malta, and Cyprus are the exceptions that do not directly regulate PRPs. For example, Ireland negotiates pharmacy markups for prescription drugs with the pharmaceutical industry.

Columns 4 and 5 present the VAT rates on prescription and OTC drugs, respectively. Column 6 shows the standard VAT rate in each country. The table shows that only Germany, Denmark and Bulgaria use the highest possible VAT bracket for RX drugs. Since catastrophic health spending—defined as a percentage of income and using a threshold of 40% of household capacity to pay for health care—is more likely among lower income households (OECD and European Union 2022, p. 177), it is likely that countries use lower VAT brackets as a mean of subsidizing poorer households. The reduced VAT brackets have long been under scrutiny, and a more uniform tax rate would likely generate more tax revenue and improve consumer welfare because the reduced VAT brackets distort relative prices between different goods and services (Mirrlees and Adam 2010). My results in Section V highlight an additional argument for the abolition of the reduced tax brackets (at least for pharmaceuticals) by showing that, in practice, the reduced tax rates can operate as a tax subsidy for manufacturers.

II Data

I use data from Fimea and the Association of Finnish Pharmacists (AFP). The first data set contains monthly package-level wholesale data on the sales value and volume of each pharmaceutical package sold on the Finnish pharmaceutical market.[7] These data measure the purchases pharmacies make from wholesalers. I complement this data set with price regulation information from the AFP. With price regulation information, I identify which products are subject to binding price caps.

For my reduced-form analysis, I restrict my estimation sample in several different steps. First, I limit my estimation window to years 2013–2014 or 24 months. Second, I include only products whose price caps remained constant throughout the period to isolate the effect of the markup reform from possible price cap renegotiations occurring at the same time. Finally, I restrict my sample so that the control group and the treatment group do not contain products within the same Anatomical Therapeutic Chemical classification

7. Due to the Finnish regulatory environment, my data avoids measurement errors arising from discrepancies between net and list drug prices caused by confidential rebates common in the U.S., particularly under Medicare Part D (Ippolito and Levy 2022).

Table 3: Reduced Form Sample Descriptive Statistics

	Treatment	Control
Mean Retail Price	63.29	140.96
	(1.22)	(20.18)
Mean Wholesale Price	42.61	107.61
	(0.93)	(16.21)
Mean Sales	18983.01	28828.84
	(464.97)	(3804.28)
# molecules	344	70
# firms	176	60
# packages	930	221
# observations	19666	4602

Notes: This table presents the descriptive statistics of the reduced form estimation sample. Standard errors in parentheses. Outcome data source: Fimea (2013–2014). Prices and sales in nominal euros.

system (ATC) group 5 (the same molecule or active ingredient). This addresses the risk of equilibrium effects, as all substitution between prescription drugs can legally occur at the molecule-strength-package size level. Thus, the substitution of drugs with different molecules could occur only by the prescribing physician's decision. This restriction also excludes a potential situation in which the products in the control group would be strategic substitutes for the products in the treatment group.

Table 3 provides the relevant descriptive statistics for this sample by treatment status. Products included in the treatment group (products with nonbinding price caps) are, on average, more expensive at the package level than products in the control group. The price difference should not be concerning, as the relative magnitudes of the markup cut under full passthrough are quite similar (4.3% vs 4.1%).[8] However, the monthly total sales are on average larger for products in the control group. The treatment group is significantly larger in size than the control group, with 344 molecules compared to the 70 molecules in the control group and more than four times the number of packages and observations. My estimation sample consists of significantly more active ingredients than the existing reduced-form literature on pharmaceuticals, with the notable exceptions of Kortelainen et al. (2023) and Granlund and Bergman (2018) whose samples consist of several hundreds of active ingredients. Other existing studies have typically studied only a few substances at a time.[9]

In my structural estimations, I complement my Finnish data sources with pharmaceutical price data from Sweden, Denmark, and Norway. I use these data to construct

8. See Table 1. Under full passthrough, the new retail prices for products with wholesale prices at the two group averages would be 58,80 and 139,38 euros.

9. For a short review, Brekke, Grasdal, and Holmås (2009) studied six molecules, Brekke, Holmas, and Straume (2011) eight and Brekke, Canta, and Straume (2015). The Pavcnik (2002) sample consists of three ATC3 classes.

Table 4: Statins Market Descriptive Statistics

	C10AA01 Simvastatin	C10AA07 Rosuvastatin	C10AA05 Atorvastatin	C10AA02 Lovastatin	C10AA03 Pravastatin	C10AA04 Fluvastatin	Total
Retail Price	27.73	58.06	35.51	23.40	28.09	24.73	36.69
	(27.51)	(44.19)	(20.96)	(13.63)	(19.28)	(14.48)	(32.13)
Package Size	73.54	63.08	68.35	77.41	73.60	69.72	69.60
	(33.82)	(35.18)	(34.60)	(32.34)	(33.88)	(34.56)	(34.61)
Unique Firms	8.84	6.26	5.76	2.92	3.55	3.07	6.43
	(1.83)	(1.35)	(0.48)	(1.18)	(0.72)	(0.66)	(2.35)
Unique Products	50.82	40.47	39.25	8.15	11.12	10.07	38.14
	(11.86)	(10.49)	(3.32)	(3.86)	(2.51)	(2.03)	(16.40)
Sales (PRP)	19.66	13.94	21.75	0.64	1.41	2.42	60.72
	(8.98)	(2.24)	(7.71)	(0.15)	(0.28)	(0.54)	(17.76)
Sales (PPP)	12.08	9.19	14.01	0.40	0.89	1.54	38.73
	(6.10)	(1.57)	(5.23)	(0.09)	(0.20)	(0.33)	(12.23)
Market Share	50.83	14.07	31.91	0.82	2.44	3.01	100.00
	(8.70)	(3.09)	(6.66)	(0.28)	(0.25)	(0.56)	(0.00)

Notes: This table presents the descriptive statistics for the variables used in the estimation. Data sample 2010–2017. Mean values are presented in the first row, and standard deviations in parentheses. Sales in annual pharmacy retail prices (PRP) and pharmacy purchase prices (PPP), both in millions of nominal euros.

Hausman-like instruments for my demand model. I describe this process in more detail in Section IV. The data sources are listed in the Appendix A. I also limit my analysis to statins (ATC4 level C10AA consisting of six molecules), and aggregate the sales data to the quarterly level. I focus on the statins market for several reasons. First, modeling all drug classes would be a burdensome exercise both conceptually and computationally. Some drug markets serve chronic diseases, while others are used for the treatment of acute diseases. With or without public reimbursement, some markets are generic markets, while others are monopoly markets with active patents, complicating the regulatory environment. Therefore, it is unlikely to find a one-size-fits-all model that fits the data well.[10] Second, the statin market has been the subject of interest in previous studies in the literature, such as Kaiser et al. (2014), which allows me to compare my estimates with the results in the existing literature.

The descriptive statistics of this sample are presented in Table 4. During the data sample, the most expensive drug on average was rosuvastatin, and the most inexpensive drug was lovastatin. Simvastatin had the largest market share at the molecule level, with an average annual market share of more than 50%. Simvastatin also had, on average, the most manufacturers and packages on the market, while lovastatin had the fewest. The average annual sales of all statins during the sample period was approximately 60 million euros at retail prices (including the VAT) or 38 million euros at wholesale prices.

III Reduced Form Evidence

The empirical design in my reduced-form analysis is based on Difference in Differences (DID). I take advantage of the fact that due to the binding wholesale price cap regulation, not all firms were able to respond to the change in the retail markups by increasing their prices. By construction, this assumes that the firms' best response is to increase their prices when the markup formulas in Table 1 were changed. This assumption is illustrated by equation (9) of my supply model in Section IV. The equation demonstrates that the partial derivatives for wholesale prices are positive with respect to the pharmacy markups and the VAT rate. My treatment group consists of products that did not have binding price caps at the time of the policy change, allowing manufacturers to increase their wholesale prices in response to the reform. In contrast, my control group consists of products that had binding price caps, which means manufacturers could not increase their wholesale prices to offset the decreased retail mark-ups.

Equation (1) presents my event study model to estimate the passthrough of the

10. Notable exceptions are (Dubois, Gandhi, and Vasserman 2022; Atal, Cuesta, and Sæthre 2022). The former studies the American and Canadian hospital drug markets, and the latter studies the Chilean retail market. Both markets lack the common regulatory environment in European countries such as Finland.

decrease in pharmacy markups:

$$y_{it} = \boldsymbol{\alpha}_i + \boldsymbol{\lambda}_t + \sum_{\tau \neq -1} \beta_\tau \text{Reform}_{i\tau} + \epsilon_{it} \tag{1}$$

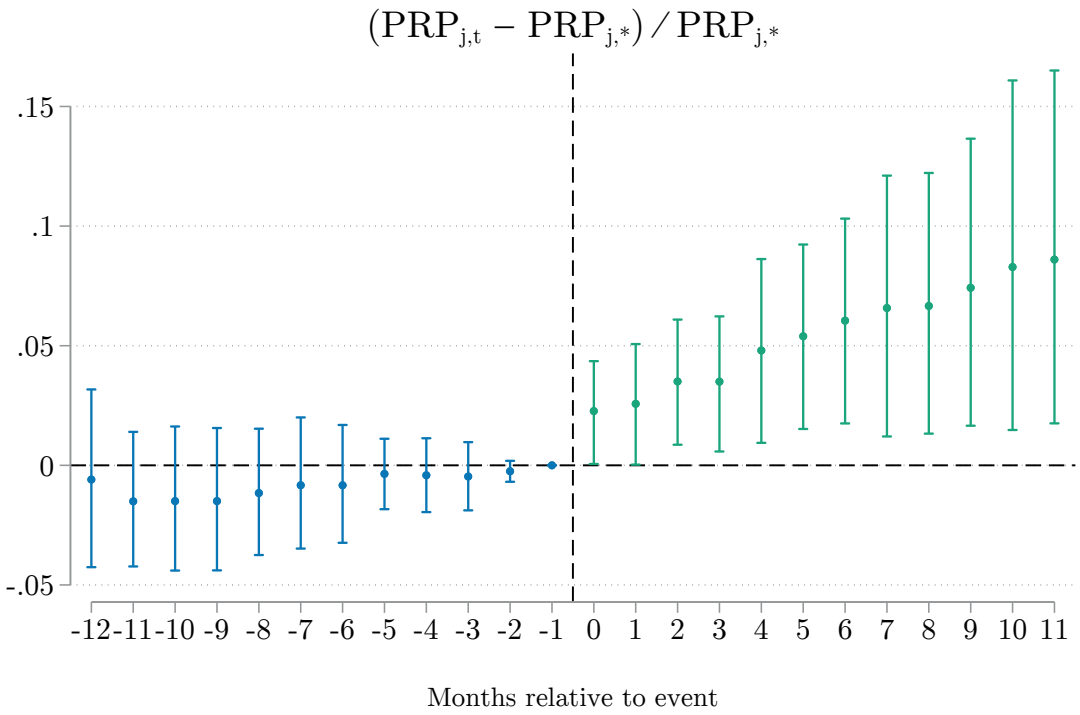
where y_{it} represents the outcome of interest, which is the percentage change in retail prices relative to the base period $t^* = -1$, the month before the regulatory change, for the product i in period t, i.e., $(PRP_{j,t} - PRP_{j,t^*})/PRP_{j,t^*}$. This outcome variable is related to the measurement of tax passthrough in the public finance literature (see e.g. Kosonen 2015; Harju, Kosonen, and Skans 2018). The vector $\boldsymbol{\alpha}_i$ denotes the package-level fixed effects and $\boldsymbol{\lambda}_t$ denotes the period (year-month) fixed effects. The $\text{Reform}_{i\tau}$ variables indicate the time-to-treatment, set to 1 for treated products at time t when τ periods have elapsed since the start of treatment. The coefficient β_τ captures the average treatment effect on the treated (ATT) from time τ to the period just before treatment. I also estimate the ATT using a canonical 2×2 DID setup where I change equation (1) by replacing $\sum_{\tau > -1} \beta_\tau \text{Reform}_{i\tau}$ with $\beta_{ATT} \text{Reform}_i$ and $\boldsymbol{\lambda}_t$ with γ_{Post}. In this case, β_{ATT} can be interpreted as the average impact of the reform in the post period.

Parameters $\hat{\beta}_\tau$ and $\hat{\beta}_{ATT}$ can be interpreted as causal under specific identifying assumptions. These assumptions are the usual: I require the parallel trends assumption (with no anticipation effects) and Stable Unit Treatment Value Assumption (SUTVA). The former assumes that the prices in the treatment group would have, on average and across all periods, evolved similarly to the prices in the control group in the absence of treatment. The latter assumes that the treatment assignment of any product does not affect the potential outcome of any other product. I discuss possible issues with respect to these assumptions at the end of this section.

The event study results are presented in Figure 1. The results during the pre-period are all precise zeros, which is supporting evidence for the parallel trends assumption. The post-period results clearly indicate that at the beginning of the new markup regulation ($t = 0$), the retail prices of the products in the treatment group started to increase relative to the control group. At the start of the new regime, the dynamic treatment effects imply that retail prices in the treatment group increase by approximately 4 percentage points relative to the control group. After 12 months, the effect is close to 10 percentage points. All dynamic effects are statistically significant at the 95% level. The ATT results are presented in Figure 2. The average effect was 0.062 and statistically significant, which corresponds to an increase of 6.2 percentage points in the retail prices of products in the treatment group.

Note that in this particular case, the products in the control group were also subject to a change in pharmacy markups. However, because these products were subject to binding price caps, pharmaceutical companies were unable to increase their prices, and mechanically the change in retail markups was fully transferred to their retail prices.

Figure 1: Event Study Estimates



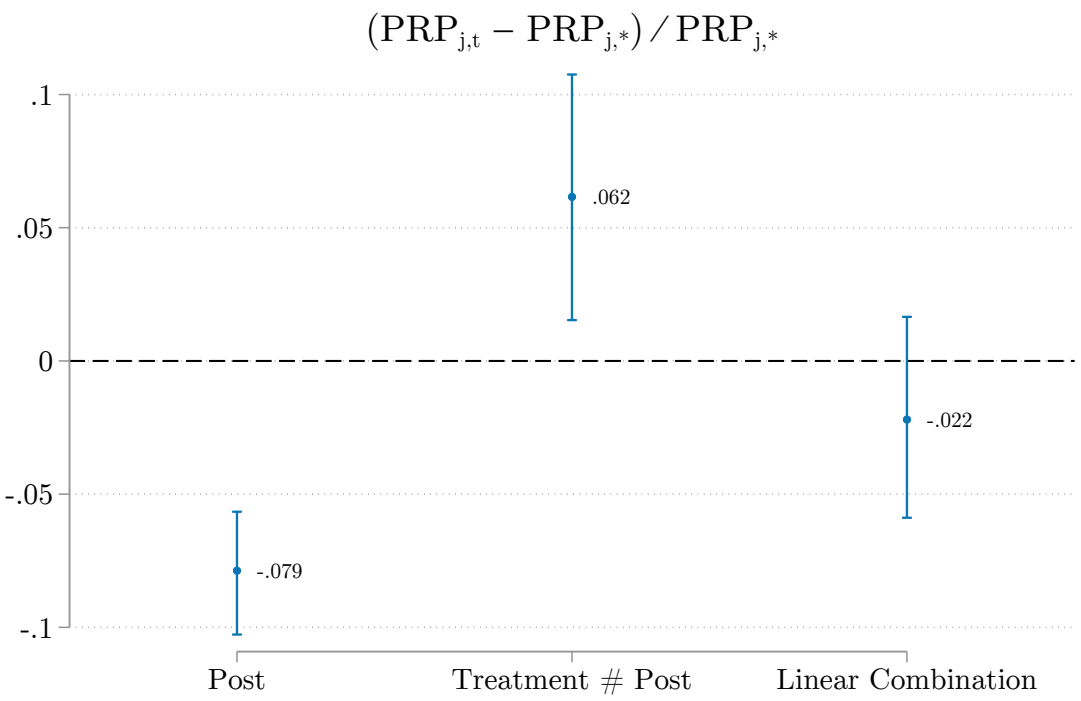
Notes: This figure present the event study estimates for equation (1). Outcome variable: Change in retail prices relative to December 2013. Estimator: Two-Way Fixed Effects (TWFE) with bootstrapped standard errors clustered at the ATC3 level (10000 replications).

Thus, the counterfactuals (the potential outcomes) in the reduced-form exercise are slightly different from those in the standard DID literature. The control group gives a counterfactual of complete passthrough to estimate the degree of partial passthrough for the products in the treatment group. To compute the actual passtrough of the markups, I compare the aggregate change of prices in the treatment group with the price changes in the control group.

Figure 2 shows the estimate of the linear combination of $\hat{\gamma}_{Post} + \hat{\beta}_{ATT} = -0.022$. However, this estimate is not statistically significantly different from zero. Since the term $\hat{\gamma}_{Post}$ is related to the case of full passthrough, I can calculate the partial passthrough rate by dividing my linear combination estimate by the term $\hat{\gamma}_{Post}$. This yields me the final estimate of 28% for the passthrough of the change in pharmacy markups. Conceptually, measuring passthrough by the comparison of consumer prices relative to full passthrough again follows the previous literature on the estimation of tax passtrough with reduced-form methods (Kosonen 2015; Harju, Kosonen, and Skans 2018).

However, my reduced form estimations suffer from possible SUTVA violations. Al-

Figure 2: ATT Estimates



Notes: This figure present the ATT estimates for equation (1). Outcome variable: Change in retail prices relative to December 2013. Estimator: TWFE with bootstrapped standard errors clustered at the ATC3 level (10000 replications). The implied passthrough can be calculated from the ATT result by dividing the linear combination with the base level. This yields $(-0.022)/(-0.079) \times 100\% \approx 28\%$.

though I restrict my sample so that there is no direct competition between the treatment and control groups, SUTVA also requires that the treatment statuses of the products within the treatment (or control group) should not affect the potential outcomes of other products in the same group. However, since they can be direct substitutes within the treatment or control group, this assumption is not satisfied (Minton and Mulligan 2024). Furthermore, it should be obvious that products that I discard from the sample could also affect these estimates.[11] The direction of the bias from SUTVA violations depends on the nature of competition between and within the control and treatment groups. However, if more products are strategic substitutes, the estimates are biased upward because the best response to any price change is in the opposite direction. If more products are strategic complements, there is a downward bias, as the best response is to adjust prices in the same direction. Empirically, the direction of the bias will depend on the distribution of substitution patterns between the products.

11. I cannot guarantee that my sample includes every product with the same molecule.

The above issues highlight how substitution patterns influence pricing dynamics. However, my results demonstrate that compared to the control group and full passthrough, products that could increase their wholesale prices had smaller decreases in retail prices after pharmacy markups were reduced.

IV A Structural Model of the Statin Market

In this section, I present my structural model for the statin market in Finland. I use the structural model to simulate counterfactual outcomes under different markup regulation and VAT systems. This is useful for three purposes. First, it allows the study of tax reforms that have not yet been implemented in the market. Second, the structural approach allows the kind of equilibrium responses from drug manufacturers that SUTVA does not in most reduced form applications. Third, it accommodates heterogeneous effects from producers.

Statins, a class of medications, are primarily used to lower blood cholesterol levels, especially bad LDL cholesterol. They function by inhibiting an enzyme in the liver essential for cholesterol production (Cholesterol Treatment Trialists Collaboration et al. 2005, 2010). Statins are divided into two generations: First-generation statins, such as lovastatin and simvastatin; and second-generation statins, including atorvastatin and rosuvastatin, which are often considered more potent. These medications are typically consumed for long periods, often for a lifetime, due to their role in the management of chronic diseases such as cardiovascular disease and the prevention of heart attacks and strokes.

I first present my discrete choice demand model, followed by the supply model and the calculation of marginal costs. I then present estimation of passthrough. I conclude the section by discussing the identification and estimation of the models.

A Demand model

Consumer i obtains indirect utility from a standard unit of dosage in package $j \in J$ in market $t \in T$ following the structure shown in equation (2). Each market t is defined as a quarter of a year, so all statin products sold in a given quarter belong to the same market across active ingredients, strengths, package sizes, and dosage forms.[12]

$$u_{ijt} = \boldsymbol{\sigma}_j^D + \boldsymbol{\tau}_t^D + \sum_k \beta^k x_{jt}^k - \alpha_i p_{jt}^r + \xi_{jt} + \epsilon_{ijt} \tag{2}$$

The terms $\boldsymbol{\sigma}_j^D$ and $\boldsymbol{\tau}_t^D$ in equation (2) represent the demand side fixed effects that include molecule, year, and quarter period dummies. The exogenous product characteristics, x_{jt}, include log package sizes and an indicator for the brand (originator drug) status. ε_{ijt}

12. I have aggregated the data to the quarterly level. Prices are calculated by dividing total sales by sold quantities.

is a consumer-specific demand shock that follows a Type I Extreme Value distribution, yielding the well-known mixed logit choice probabilities (Berry and Haile 2021).

I set the market share of the outside option to 5%. The assumption corresponds to defining the market potential of the statin market. This value is taken from Kari et al. (2024), who report the observed share of non-dispensed simvastatin prescriptions in Finland during 2020–2022. Although the estimate does not capture cross-market variation, it is based on actual prescription data and is one of the most accurate measures of market potential in the pharmaceutical demand literature using aggregate data.

The endogenous term in equation (2) is the price term, p^r_{jt} which is assumed to be correlated with the unobserved quality or demand shock ξ_{jt}. This is due to the dependence between profit-maximizing prices and other unobservable factors that also affect demand. For example, firms are likely to increase their prices if there is a positive demand shock. The main endogeneity issue arises from time- and product-specific shocks not controlled by fixed effects. In the statin market, high consumer inertia means that patients tend to stick to their initial prescription, leading older patients to use first-generation statins while newer statins are prescribed to younger or higher-risk patients. These differences in consumer bases, shaped by evolving demographics or regulatory trends such as reimbursement rules, can generate demand shocks that fixed effects cannot fully capture.

Because drugs are sold in different strengths and package sizes, the price term cannot be measured at the actual package-level price. Otherwise, the econometrician would impose a greater weight on small packages with less potent drugs in the calculation of market shares. The existing literature has solved this problem by measuring quantities and prices in terms of some standard units (Dubois and Lasio 2018; Dubois and Sæthre 2020), or using expenditure shares rather than quantity market shares. The latter approach was first used by Björnerstedt and Verboven (2016), and has since been used in the structural pharma-literature by Atal, Cuesta, and Sæthre (2022). In that specification, prices would enter logarithmically and the market shares and the size of the potential market are measured in expenditure shares. I choose to follow the former approach, which is closer to the canonical demand models in the literature. I measure the price by the price of the package divided by the number of Defined Daily Dosagess (DDDs) included in a package.[13] Therefore, market shares will also be measured in terms of the number of doses sold.

An important part of my demand model is the distributional assumption on the price coefficient α_i. In a standard multinomial logit case, the model would assume that consumers have homogeneous preferences over price. However, this assumption has stark consequences on the elasticities and rates of passthrough that the underlying demand system can support. Miravete, Seim, and Thurk (2023) show that a standard multinomial

13. DDD is a measure used in pharmaceutical and health studies. It was developed by the World Health Organization (WHO). The DDD is defined as the normal maintenance dose per day for a drug used for its main indication in adults.

logit model can only produce standard passthrough estimates truncated at 100%. In comparison, allowing heterogeneity in the price term allows for more flexible elasticities and rates of passthrough. In my main approach, I follow Miravete, Seim, and Thurk (2023) by assuming that the price coefficient follows a log-normal distribution, $\alpha_i \sim \log \mathcal{N}(\alpha, \Sigma_\alpha)$, which imposes downwards-sloping demand for all consumers. For comparison, I also consider a specification where the coefficient follows a normal distribution.

Note that studying drug markets using a discrete choice model of demand is inherently different from studying the demand for breakfast cereals or cars. Under generic substitution, consumers can only freely choose between the exact prescribed product and its direct substitutes. For example, a patient with a prescription for a dose of 10 milligrams of simvastatin can only substitute between exactly those products, not 20 milligrams of simvastatin or any other statin. Therefore, in most cases, consumers are tied to the decision that the prescribing physician makes. As argued by Crawford and Shum (2005) and Dubois and Lasio (2018), drug demand with aggregate data is always a mixture of physician prescriptions, regulation, and patient preferences, not just pure consumer choice. My model does not separate interactions between patients, physicians, and pharmacists during prescription and purchase. For these reasons, I also abstract away from explicitly modeling regulation related to consumer choice—such as reference pricing—in my demand model.

B Supply model

Although my demand model does not include regulation that influences consumer choice, I explicitly model the vertical structure and markup regulation in my supply model. This approach follows the the vertical structure in the supply model in Miravete, Seim, and Thurk (2018, 2020). However, the maximum wholesale price regulation makes the estimation of the supply side slightly more difficult than in standard IO applications. I assume that firms compete Bertrand-Nash, that is, firms maximize:

$$\begin{aligned} \underset{p^w_{jt}}{\text{maximize}} \quad & \sum_{j \in J^f} \overbrace{(p^w_{jt} - c_{jt})}^{\text{Wholesale markup}} \times M_t \times s_{jt}(\underbrace{p^r_{jt}(p^w_{jt})}_{\substack{\text{Markup} \\ \text{formula}}}, \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}), \\ \text{subject to} \quad & p^w_{jt} \leq \bar{p}^w_{jt} \end{aligned} \tag{3}$$

with respect to their wholesale prices p^w_{jt}. Because of regulation, firms are subject to price caps $\bar{p}^w_{jt}$ imposed by (or negotiated with) the government. The vertical market structure is well visible in equation (3): markups for pharmaceutical manufacturers consists of the

price difference between the wholesale price and the marginal cost c_{jt}, but the market shares s_{jt} are a function of retail prices p^r_{jt} which is a function of wholesale prices. Note that the marginal costs, as well as the prices, are in terms of DDDs. The objective function yields the following first-order conditions:

$$s_{jt}(\mathbf{p}^r(\mathbf{p}^w)), \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}) + \sum_{m \in J^f} (p^w_{mt} - c_{mt}) \times s_{mt}(\mathbf{p}^r(\mathbf{p}^w)), \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}) \times \frac{\partial s_{mt}}{\partial p^w_{jt}} \geq 0 \tag{4}$$

where the last term captures the changes in quantities (of all products of firm j) when the wholesale prices p^w_{jt} change. When the price caps do not bind, the following first-order condition (Equation 5) holds:

$$\sum_{m \in J^f} (p^w_{mt} - c_{mt}) \times s_{mt}(\mathbf{p}^r(\mathbf{p}^w)), \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}) \times \frac{\partial s_{mt}}{\partial p^w_{jt}} + s_{jt}(\mathbf{p}^r(\mathbf{p}^w)), \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}) = 0 \quad \text{for} \quad p^w_j < \bar{p}^w_j. \tag{5}$$

The first-order condition for products with binding price caps is given in equation (6):

$$\sum_{m \in J^f} (p^w_{mt} - c_{mt}) \times s_{mt}(\mathbf{p}^r(\mathbf{p}^w)), \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}) \times \frac{\partial s_{mt}}{\partial p^w_{jt}} + s_{jt}(\mathbf{p}^r(\mathbf{p}^w)), \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}) \geq 0 \quad \text{for} \quad p^w_j = \bar{p}^w_j. \tag{6}$$

Note that marginal costs cannot be backed out for products with a binding constraint ($p^w_j = \bar{p}^w_j$). In most cases, the calculation of counterfactual prices requires estimates for marginal costs. For the current exercise, this is not a major concern, because most products are priced below their price ceilings. In the case of a retail markup or a tax decrease, the best response for firms is to increase the wholesale prices for these products. However, if prices were strategic substitutes, it might be profitable for some of the firms to lower the prices of the products with binding price caps in response to their competitors' price increases. A BLP demand system with Bertrand-Nash pricing allows, in theory, that prices can be either strategic substitutes or complements. If all prices are strategic compliments, the best response to competitors' price increases (cuts) is a price increase (cut).

In the existing literature, Dubois and Lasio (2018) estimate manufacturer markups under maximum price regulation in the anti-ulcer market in France. In their model, they estimate the marginal costs using data from unregulated markets in Germany and the United States. The approach in Dubois and Lasio (2018) does not require that the econometrician observes which products have binding price caps, but their approach

does require that marginal cost estimates for every product in unconstrained markets are available. Fan and Zhang (2022) take a similar approach in an application to the cell phone markets in China. Using data from markets without price ceilings, they project their marginal costs on observable firm characteristics and estimate the empirical distribution of supply shocks. They then simulate and solve for the expected marginal costs for the products with binding price ceilings in regulated markets.

Unfortunately, I cannot follow either approach, even though I have data from four Nordic countries. The Dubois and Lasio (2018) approach is not feasible because only a subset of products sold in Finland is available in other Nordic countries. Furthermore, the Fan and Zhang (2022) approach is not appropriate because all Nordic countries either directly regulate pharmaceutical price ceilings or negotiate price caps with pharmaceutical manufacturers.

I overcome this limitation by using inputed marginal costs. To be more precise, I first regress the Bertrand-Nash marginal costs from equation (5) on observable firm and product characteristics. The regression model is presented in equation (7):

$$\log c_{jt} = \boldsymbol{\gamma}\mathbf{Y}_{jt} + \omega_{jt}. \tag{7}$$

where $\mathbf{Y}_{jt}$ consists of a constant, log package size, at set of ATC5, brand status and firm dummies and market fixed effects. The ω_{jt} represent the unobserved supply shocks that affect marginal costs. I estimate the marginal cost parameters $\hat{\boldsymbol{\gamma}}$ with OLS and use these estimates to predict marginal costs $\hat{c}_{jt}$ for the products with binding price caps (equation 6). This means that I assume that the supply shocks ω_{jt} are mean zero and are orthogonal to $\mathbf{Y}_{jt}$ in the model. The difference from the approach Fan and Zhang (2022) is that I cannot estimate an unconditional empirical distribution of the shocks to simulate the expected marginal costs.

C Passthrough and markups

I estimate the passthrough of retail markups and value-added taxes by computing counterfactual prices under different markup and tax regimes. From equation (4), I obtain the following expression:

$$\mathbf{p^w} = \mathbf{c} - [\underbrace{\boldsymbol{\Omega}}_{\substack{\text{Ownership} \\ \text{matrix}}} \times \overbrace{\boldsymbol{\Delta}^{w'}}^{\frac{\partial s_m}{\partial p_j^w} \times \frac{\partial p^r}{\partial p^w}}]^{-1} \times s_{jt}(\mathbf{p}^r(\mathbf{p}^w)), \mathbf{x}, \boldsymbol{\xi}, \boldsymbol{\theta}). \tag{8}$$

Because the government sets the margins, the $\boldsymbol{\Delta}^w$ term in equation (8), representing the demand derivatives with respect to wholesale prices, can be expressed as

$$\boldsymbol{\Delta}^w = \boldsymbol{\Delta}^d \boldsymbol{\Delta}^p = \underbrace{\begin{bmatrix} \frac{\partial s_1}{\partial p_1^r} & \cdots & \frac{\partial s_1}{\partial p_J^r} \\ \vdots & \ddots & \vdots \\ \frac{\partial s_J}{\partial p_1^r} & \cdots & \frac{\partial s_J}{\partial p_J^r} \end{bmatrix}}_{\text{Demand Jacobians w.r.t price}} \times \begin{bmatrix} \frac{\partial p_1^r}{\partial p_1^w} & \cdots & \frac{\partial p_J^r}{\partial p_J^w} \\ \vdots & \ddots & \vdots \\ \frac{\partial p_J^r}{\partial p_1^w} & \cdots & \frac{\partial p_J^r}{\partial p_J^w} \end{bmatrix} \tag{9}$$

where all elements in the second matrix are known from the regulatory rules. The cross-derivates $(\partial p_{j'}^r)/(\partial p_j^w) = 0$, and the diagonal elements $(\partial p_j^r)/(\partial p_j^w)$ consist of the retail markup and the VAT. I know the markup function and the VAT rate:

$$\frac{\partial p_j^r}{\partial p_j^w} = \begin{cases} \rho_1 + \tau & \text{if } p_{jt}^w \times \text{DDDs per package} \leq 9.26 \\ \rho_2 + \tau & \text{if } 9.26 \leq p_{jt}^w \times \text{DDDs per package} \leq 46.25 \\ \rho_3 + \tau & \text{if } 46.26 \leq p_{jt}^w \times \text{DDDs per package} \leq 100.91 \\ \rho_4 + \tau & \text{if } 100.92 \leq p_{jt}^w \times \text{DDDs per package} \leq 420.47 \\ \rho_5 + \tau & \text{if } p_{jt}^w \times \text{DDDs per package} > 420.47 \end{cases} \tag{10}$$

where ρ's are approximations of the derivative of the piece-wise markup function with respect to the wholesale price and $\tau = 10\%$ is the VAT rate for pharmaceuticals. The estimation of the passthrough elasticities of the markups and taxes for product j follows after a counterfactual simulation by calculating:

$$\psi_j = \frac{\overbrace{p_j^{New} - p_j^{Old}}^{\text{Price difference}}}{\underbrace{\Delta\left(\rho + \tau\right)}_{\text{Cost shock}} \times p_j^{Old}} \tag{11}$$

where the denominator denotes the price level under full passthrough. Thus, equation (11) compares the change in retail prices under two markup or tax regimes (new and old) and compares it to a case of full passthrough, that is, how much prices would have changed had the change in the tax rate been transferred to retail prices one-to-one.

D Identification

I need instruments to identify the parameters related to price sensitivity, namely α and σ_α. I utilize two types of instruments. First, I construct Hausman-like instruments using price data from other Nordic countries. These instruments can be classified as "cost shifters", as they capture common shocks to supply across countries and markets. The second type of instruments consists of Gandhi and Houde (2020) differentiation instruments constructed from one of the exogenous product characteristics, the package size. This instrument is a so-called "demand shifter", as it measures aggregate changes in the characteristics of competing products, which shifts the demand of the product

in question. I do not construct any differentiation or BLP instruments from the other exogenous product characteristics, all of which are dummies. In practice, they would capture aggregate changes in the consumers' choice sets from product entry and exit, and in many cases they would be collinear with some of my fixed effects. Therefore, in total, I have four instruments in my use: Three Hausman-like instruments, and one differentiation instrument.

Ideally, my Hausman instruments would consist of the prices of the same products in other Nordic countries. However, most products are not sold in other countries, let alone in all of them. The typical case in the literature is to use simple molecule-level averages (Atal, Cuesta, and Sæthre 2022), but the variation provided by such an aggregated measure is limited, especially if product and market fixed effects are included in the model. Therefore, I take a slightly more sophisticated approach that resembles the process of imputing marginal costs in Section B. First, for packages that are sold in a neighboring country, I use its own price in the neighboring country as the value of the instrument. For the other packages, I rely on an imputation approach. To be more precise, I estimate the following hedonic regression where the dependent variable, log wholesale price, is regressed on a set of observables X_{jt}:

$$\log p_{jt}^{w} = \boldsymbol{\beta}\mathbf{X}_{jt} + \varepsilon_{jt} \tag{12}$$

where $\mathbf{X}_{jt}$ comprises a constant, package sizes measured in DDDs and a set of molecule, firm, time period, branded, and reimbursement dummies. I estimate the model for Sweden, Denmark and Norway, and use the estimated country-specific estimates $\hat{\boldsymbol{\beta}}$ with Finnish data to predict the log prices for the products that were not on the market in the other countries. Finally, I use the exponentiated predictions as instrument values.[14] In practice, these predicted values represent a type of conditional mean for the prices of similar products sold in the neighboring country.

My approach is similar in spirit to Barahona, Otero, and Otero (2023) who study the Chilean market for breakfast cereals. They construct simulated instruments by regressing cereal prices to known input prices and fixed costs. Barahona, Otero, and Otero (2023) then use the predicted prices as instruments for the actual prices. Compared to their instruments, my instruments are a combination of actual Hausman instruments and simulated instruments.

E Estimation

Since my supply model does not directly yield marginal costs for all products, I estimate the demand and supply models separately. I estimate the demand side with the suggested

14. Using logs in estimation and converting them to levels is done due to significant outliers in pharmaceutical prices across molecules.

best practices from Conlon and Gortmaker (2020). The model is first estimated using the instruments described in Section D. Then, I estimate the Chamberlain (1987) and Reynaert and Verboven (2014) optimal instruments and solve the demand model again with them. These are the final results reported in Section V. I use 1000 Halton draws to simulate the agents used in the integration over the individual choice probabilities.

I use the demand-side estimates to compute the marginal costs for products without binding price caps (equation 5). I use these marginal cost estimates to compute counterfactual prices by changing the diagonal entries inside the matrix $\mathbf{\Delta}^p$ in equation (9). However, for these counterfactual prices, I impose the simplifying assumption that the demand jacobians $\mathbf{\Delta}^d$ in equation (9) are fixed. With this assumption, I avoid solving the equilibrium prices from the full model. Therefore, my results should be interpreted only in the sense of a partial counterfactual simulation.

V Results

A Demand and Supply Estimation Results

The estimates of my structural model are presented in Table 5. My main specification is the log-normal random-coefficients model, but I also present the results from a simpler multinomial logit model and a standard random-coefficient model for comparison. In all of the models, the coefficient for log package size is positive, meaning that, on average, consumers prefer larger package sizes over smaller ones. The price coefficients for the multinomial logit model and the standard random coefficient model are both negative. Note that the price coefficient for the log-normal model can be positive or negative as α_i is always strictly positive and is applied to negative prices during the estimation. Unlike in the standard random coefficient model, this ensures that demand is downward sloping for all consumers.

The multinomial logit model yields a mean own price elasticity of -0.95 and, most importantly, a mean marginal cost of -10 euro cents per DDD. The negative marginal costs are a clear indicator that the multinomial logit model yields price elasticities that are too small. The distribution of the marginal costs of the multinomial logit model is even more concerning: 61% products have negative marginal costs. The other two models perform significantly better and imply positive marginal costs, meaning that modeling heterogeneity in consumer price sensitivity significantly improves consistency between the demand and supply models.

The two random-coefficient models give very similar mean elasticities and marginal costs, but the underlying elasticity distributions differ significantly. The mean own-price elasticity of the log-normal model is -2.04 against -1.99 of the canonical BLP-model. The estimated marginal costs of both models are, on average, both positive, with the

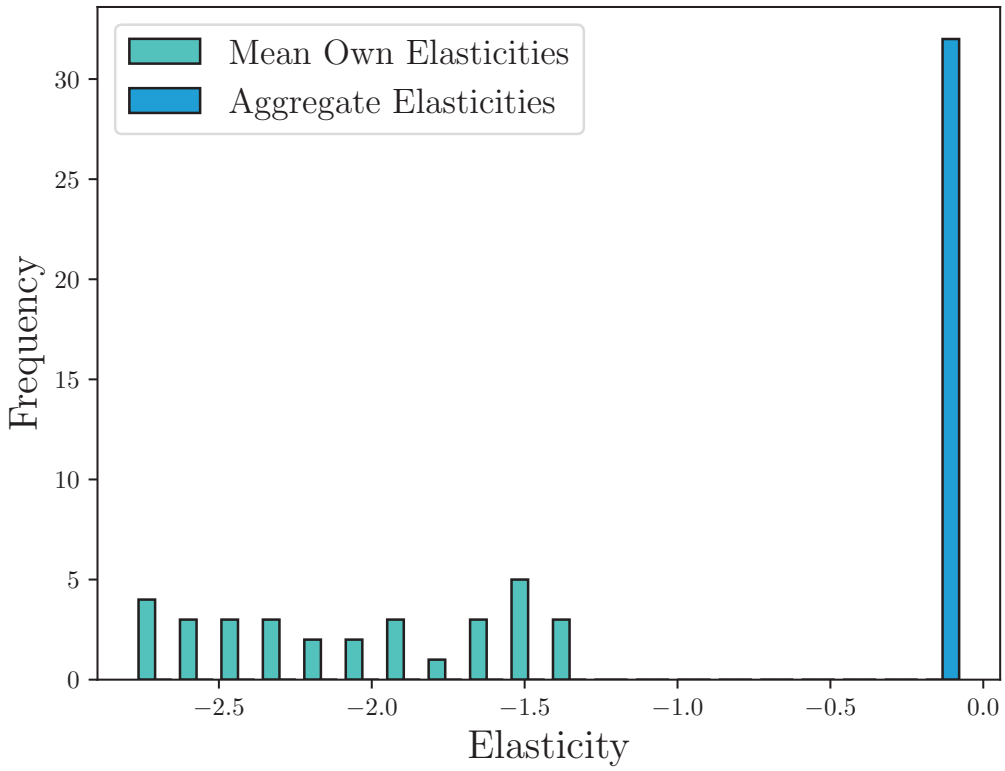
Table 5: Demand Model Results

Parameter	Logit	Random Coefficients Log-Normal	Random Coefficients Normal
	Panel A: Linear Parameters		
α Prices	-1.82	1.88	-5.81
	(0.17)	(0.35)	(0.77)
β Log Package Size	1.63	1.18	1.16
	(0.06)	(0.14)	(0.11)
Molecule dummies	Yes	Yes	Yes
Year-Quarter dummies	Yes	Yes	Yes
Quarter dummies	Yes	Yes	Yes
	Panel B: Nonlinear Parameters		
σ Prices		1.03	2.45
		(0.31)	(0.33)
	Panel C: Additional Statistics		
Mean Elasticities	-0.95	-2.04	-1.99
	(0.03)	(0.08)	(0.06)
Mean Costs	-0.10	0.13	0.16
	(0.00)	(0.00)	(0.01)
Min Costs	-0.47	-0.05	-1.51
Share of Negative Costs	0.61	0.05	0.11
Mean Passthrough	-0.40	0.58	0.52
	(0.02)	(0.00)	(0.00)

Notes: This table presents the demand model estimates. Standard errors are in parentheses. Panel A presents the estimates on the linear parameters of the demand model, representing the mean tastes of the consumers. Panel B presents the non-linear parameters, representing the standard deviations of consumer tastes with respect to prices. Panel C presents the mean own prices elasticities of demand, marginal costsand the rates of passthrough for all products.

log-normal model producing an average marginal cost of 13 euro cents per DDD. Both models produce some negative marginal costs. For the log-normal random coefficients model, 5% of all marginal costs are below zero, and for the standard random coefficients model, negative marginal costs account for 11% of the products. I present the empirical distribution of the elasticity and marginal cost estimates for my preferred model—the log normal random coeffients model—in Figures 3 and 4. Even though the negative marginal costs are a clear minority, they can be seen as a concerning result, suggesting that even the log normal random coefficients model cannot produce large enough elasticities to rationalize the observed prices under Bertrand-Nash pricing. However, my elasticity estimates are in line with those in the existing literature. Using a similar demand model

Figure 3: Own and Aggregate Elasticities



Notes: This figure presents the own price elasticity of demand of the log-normal random coefficient model of statin demand.

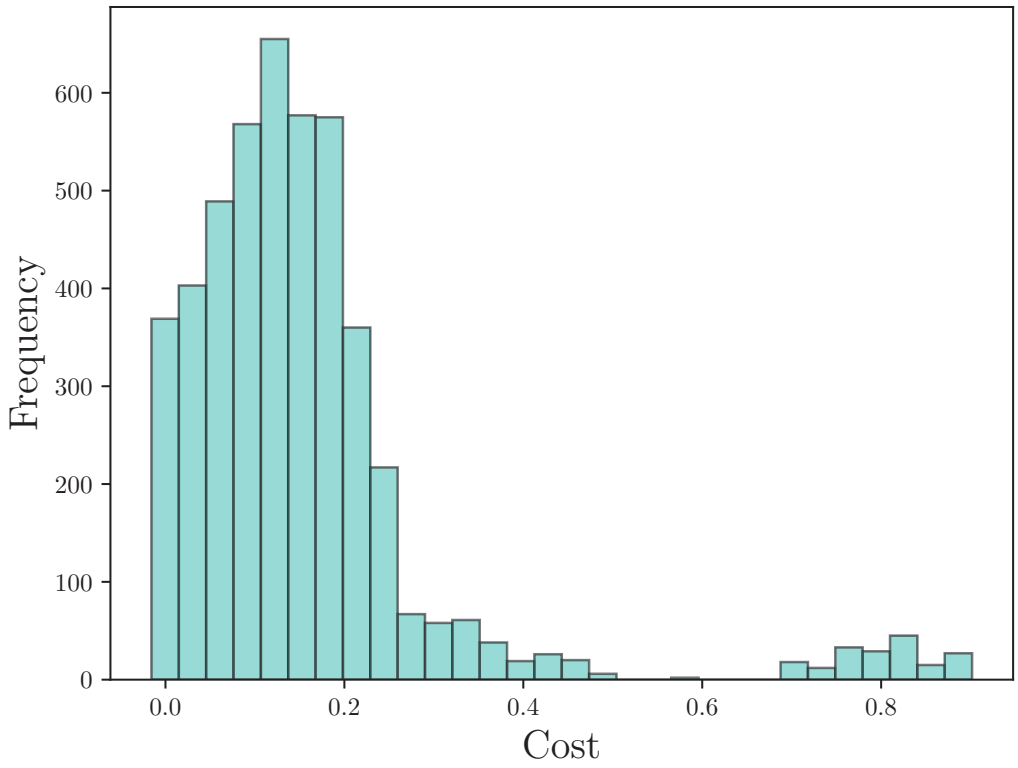
Kaiser et al. (2014) estimated a median own price elasticity of 2.52 in the Danish statin market.

The mean own elasticities are in the range $[-3, -1.5]$ and are significantly higher in magnitude than the mean aggregate elasticities, which are concentrated below -0.5. The small and concentrated aggregate elasticities imply that estimated aggregate demand is almost completely inelastic and hardly varies between markets (quarters). The inelastic aggregate demand aligns well with the parameterization of the outside share, based on the observed rate of unfilled prescriptions in the statin market.

The mean passthrough rate—implied by the simulated prices and equation (11)—is 58% for the log-normal coefficient model and 52% for the normally distributed coefficient model.[15] This means that more than half of the retail markup decreases in 2014 were carried forward to consumer prices. The rest was taken up by the manufacturers, who

15. The mean passthrough rate of the logit model was negative 40%, further indicating its poor performance. The logit model cannot capture heterogeneity in consumer price sensitivity, which significantly restricts its implied substitution patterns.

Figure 4: Distribution of Marginal Costs



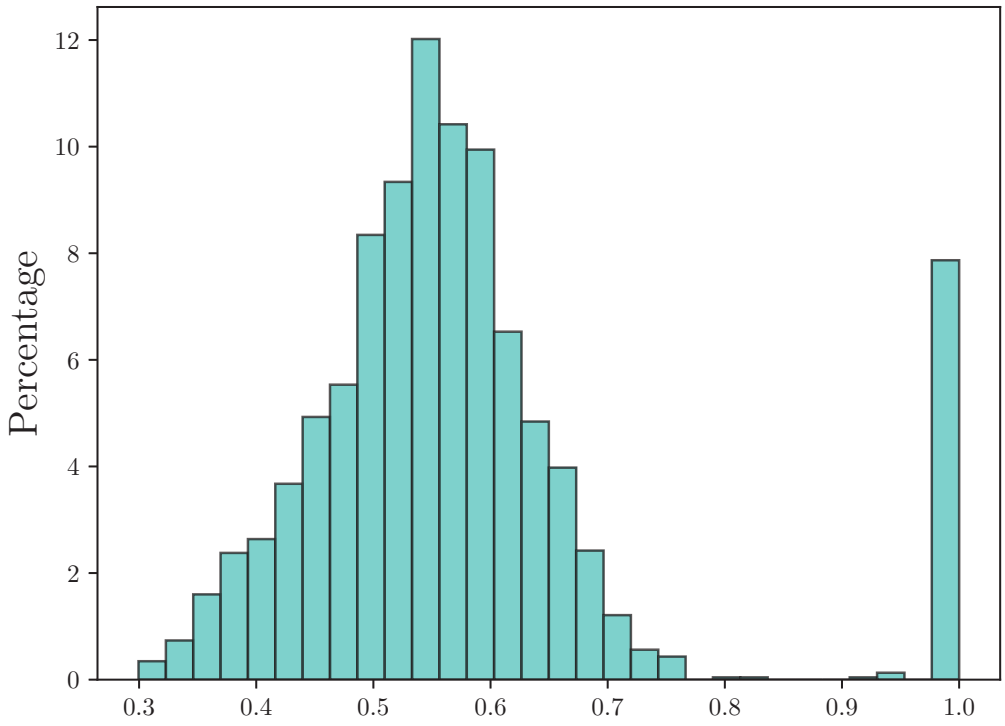
Notes: This figure presents the Bertrand-Nash marginal costs from log-normal random coefficient model of statin demand.

increased their wholesale prices.

The estimate is significantly higher than my reduced-form estimate of 28%, but because the sample consists of different drug markets, they cannot be directly compared. I present the distribution of passthrough rates for my preferred model in Figure 5. The share of the products with binding price caps is visible in the graph as the mass just below one, representing full passtrough. For the other products with non-binding price caps, the distribution appears to be approximately normally distributed, suggesting a symmetrical spread of passthrough rates centered around the peak between the 50–60% rate range.

In Table 6, I decompose the passthrough rates from the log-normal model by regressing the log passthrough rate on the exogeous product characteristics—molecule dummies and an originator-brand dummy—and the markups implied by my supply model. The markup variable is particularly important because both theoretical and empirical results from the literature suggest that firm market power has a diminishing effect on passthrough rates (Miravete, Seim, and Thurk 2018; Genakos and Pagliero 2022; Galloway and Li 2023).

Figure 5: Distribution of Passthrough Rates



Notes: This figure presents the passthrough distribution of the log-normal random coefficient model of statin demand.

The first column presents the results for the full sample of products, and the second column presents the results for products without binding price caps, i.e. products whose marginal costs have not been imputed. Both models indicate that products with larger markups have lower passthrough rates on average. This implies directly, as in Miravete, Seim, and Thurk (2018), that firms with greater market power absorb a larger share of the tax incidence. This finding is further supported by branded products having lower passthrough rates, suggesting that manufacturers of branded products possess greater market power than generic manufacturers. For the first sample, the findings suggest that the passthrough rates are on average 5–20 percentage points larger for the other molecules than for simvastatin (base level). The differences in passthrough rates between statin molecules can be further analyzed in the context of Table 4. During the sample period, simvastatin products had a combined market share of more than 50%. Atorvastatin, which had the second highest total market share of 30%, had on average five percentage points higher passthrough rates than simvastatin. For the third largest statin, rosuvastatin, with

Table 6: Passthrough Results

Dependent Variable:	Log(Passthrough)	Log(Passthrough)
Sample:	Full	No binding price caps
Model:	(1)	(2)
Variables		
Lovastatin C10AA01	0.0444*	-0.0200
	(0.0184)	(0.0144)
Pravastatin C10AA03	0.0216	-0.0165
	(0.0139)	(0.0105)
Fluvastatin C10AA04	0.0966***	-0.0047
	(0.0148)	(0.0137)
Atorvastatin C10AA05	0.0396***	-0.0135*
	(0.0082)	(0.0065)
Rosuvastatin C10AA07	0.1402***	0.0306***
	(0.0081)	(0.0067)
Branded	-0.1488***	-0.0640***
	(0.0083)	(0.0087)
Markup	-0.4691***	-0.1537***
	(0.0111)	(0.0117)
Fixed-effects		
Year-Quarter	Yes	Yes
Observations	2313	1548
R^2	0.2916	0.4007

Notes: This table presents the results from a regression model where the dependent variable is the logarithm of the estimated passthrough from Equation (11). The sample consists of the yeaars 2014–2017. The first column uses the whole data and the second column uses only data on products without binding price caps and whose marginal costs are computed directly from Equation (5).

a 14% total market share, the passthrough rates were on average 20 percentage points higher than those for simvastatin.

When conditioning on non-binding price caps, as in Column 2, the coefficients for the molecule dummies decrease substantially and lose statistical significance, except for rosuvastatin. The market power proxies—markups and the brand dummy—retain their negative signs but diminish in magnitude. This is possibly explained by the fact that the sample excludes firms with higher prices (binding price caps), and possibly greater markups.

B Counterfactual Simulations

I expand the analysis by simulating changes in revenues, manufacturer profits, consumer expenditure, VAT revenue, pharmacy profits, and quantities sold in in terms of DDDs

Table 7: Revenue Effects of the 2014 Reform and Subsequent VAT Changes

	Base	Absolute Change	Relative Change (%)
	Panel A: 2014 reform		
Quantity	767875.12	2106.80	0.27
Revenue	184568.45	6064.55	3.29
Profits	84551.05	4939.92	5.84
Expenditure	296385.70	-5373.04	-1.81
VAT Revenue	26944.15	-488.46	-1.81
Pharmacy Profits	77157.35	-9953.76	-12.90
Prices	36.55	-0.74	-2.03
	Panel B: 2014 reform + 14% VAT		
Quantity	767875.12	598.91	0.08
Revenue	184568.45	1802.66	0.98
Profits	84551.05	1517.39	1.79
Expenditure	296385.70	-1529.90	-0.52
VAT Revenue	26944.15	9266.21	34.39
Pharmacy Profits	77157.35	-10345.50	-13.41
Prices	36.55	-0.12	-0.33
	Panel C: 2014 reform + 24% VAT		
Quantity	767875.12	-3224.68	-0.42
Revenue	184568.45	-7762.02	-4.21
Profits	84551.05	-6089.72	-7.20
Expenditure	296385.70	7882.22	2.66
VAT Revenue	26944.15	31946.41	118.57
Pharmacy Profits	77157.35	-21858.22	-28.33
Prices	36.55	1.36	3.74

Notes: This table presents the aggregate changes in sold quantities, manufacturer revenues, manufacturer profits, consumer expenditure, VAT revenue and pharmacy profits between 2014–2017. Panel A presents the results for the actual 2014 markup change. Panel B presents the results for the 2014 reform with a VAT increase from 10% to 14%. Panel C presents the results for a VAT increase from 10% to 24%. Fixed demand jacobians are assumed. The first column presents the base case of the former markup regime. The absolute changes are in thousands of units and the relateive changes are in percentage points.

when pharmacy markups and the VAT rate are changed. Because my demand model abstracts away from regulation and reimbursements, the expenditure measurement is the sum of consumer copayments and government reimbursements. This framework does not fully account for potential welfare losses that arise from consumers reducing consumption; however, such losses are mitigated by the public insurance system, which covers a significant portion of pharmaceutical expenditures. In theory, these adverse effects could be offset with appropriate transfers.

The purpose of these counterfactual simulations is to demonstrate that part of the

incidence of both policies falls on the supply side. First, this increases the social cost of using reduced VAT rates to subsidize demand. Second, it shows that government-imposed retail markups have a similar incidence and that lowering them to generate savings ends up benefiting pharmaceutical manufacturers.

I specify three different specifications. First, I estimate the effects of the 2014 reform, where pharmacy markups decreased. Second, I also conduct two other counterfactual simulations, where I demonstrate the effects of the VAT as policy tool in comparison to the regulated pharmacy markups. In these counterfactuals, I augment the markup decrease by increasing the VAT rate to either 10% or to 24% from 2014 and onward. In practice, this corresponds to a scenario in which the social planner cuts the markups but compensates for this by increasing the VAT rate for pharmaceuticals. The counterfactual VAT rates are indirectly imposed by EU regulation: Member states are allowed to have up to the three VAT brackets, and in Finland these categories were 10%, 14% and 24% between 2013–2024.[16]

I present these findings in Table 7. These calculations are aggregated over the years, but I present the annual absolute and relative changes in Appendix Tables A.2 and A.3. Starting from Panel A in Table 7, aggregate consumer savings during 2014–2017 were 5.3 million euros or 1.81% relative to the old markup system. Manufacturers increased their revenues annually by approximately six million euros (3%), and their profits increased by five million euros (6%). Quantities sold increased by roughly 0.27%. Pharmacy profits decreased by approximately ten million euros combined (13%). Because consumer prices decreased and the effects on quantities were minimal, VAT revenues decreased by approximately 0.4 million in total (2%). The results from Panel A imply that roughly half of the changes in pharmacy revenues benefited consumers (and the reimbursement system), while the other half was captured by pharmaceutical manufacturers. These changes are mainly in line with the average passthrough rates (Table 5), and the minor difference is explained by the differences in the prices and market shares of products. These calculations demonstrate how ad valorem retail price regulation in the pharmaceutical sector is passed through the supply chain: The decrease in pharmacy markups in 2014 led to increases in wholesale prices—and by extension—mostly import costs of drugs.

Panel B presents the results for the smaller VAT increase and Panel C presents the results for the larger tax increase. The contrast to the estimates in Panel A is stark: For the smaller tax increase, manufacturer revenues and profits increase two-thirds less, and for the larger tax increase, the aggregate revenues and profits decrease by 7.8 million (4%) and 6.1 million euros (7%) in total. Panel B also shows that consumer expenditure decreases by roughly 1.5 million euros (0.5%), while tax revenues increase between almost ten million euros (34%). For the larger tax increase, VAT revenue increases significantly more, raising 32 million euros more (119%). However, consumer expenditure increases by

16. Not including the zero VAT rate.

7.8 million euros (3%). Pharmacy profits fall by ten million euros (13%) in the case of the smaller tax increase and by 21 million euros (28%) in the case of the larger tax increase.

The main difference between the results in Panel B and C is the change in manufacturer revenues and quantities. In Panel B, consumer prices still decrease relative to the pre-2014 regime, which is visible from the increase in total quantities sold. Wholesale prices and profits also increase, as does government tax revenue. Because pharmacies lose, Panel B represents a transfer of rents from pharmacists to manufacturers, consumers, and the government. In Panel C, manufacturers no longer profit from the change in the regulatory regime. Consumers also lose, as their expenditure increases and statin consumption decreases relative to the old regime.

The results on consumer welfare should be interpreted in the context where the government pays roughly 70% of all prescription drug expenditure. Using this number as a benchmark for the numbers in Panel C, it can be approximated that aggregate consumer copayments increased by 2.3 million euros and government tax revenues net reimbursement costs increased by 26.4 million euros. Although this is only a crude calculation, it shows that within a generous tax-funded reimbursement system, higher drug costs due to VAT are not necessarily a concern. To address consumer welfare losses from reduced demand, the government can offset the burden of higher drug prices with other transfers or adjust the copayment caps within the reimbursement system. In this context, the usual policy of small or even zero VAT rates in EU countries is ill advised.

My estimation results consider only the statin market, and although the external validity on the passthrough of pharmacy markups might carry on to other drug markets, they are not equilibrium calculations. First, they do not consider endogenous entry with respect to pharmaceutical manufacturers or the profitability of pharmacies. Both are important for the counterfactual presented in Panel C in Table 7. A decrease in pharmacy profitability could prompt product exits, reducing consumer welfare directly through lower product variety and indirectly through higher prices resulting from decreased competition. Additionally, lower pharmacy profitability could result in retailer exits, potentially increasing consumers' travel times to pharmacies.

Second, although the public reimbursement system covers 70% of all prescription drug expenditures, there is significant heterogeneity between markets and consumers. For example, consumer heterogeneity can arise through the reimbursement system, as some consumers become eligible for full reimbursement after exceeding the annual copayment cap. On the other hand, market heterogeneity is illustrated by large markets without public reimbursement, such as contraceptives. In these markets, an increase in consumer prices would result in welfare losses fully borne by a single demographic: women of reproductive age.

However, my results suggest that removing tax subsidies can finance redistribution by capturing rents from pharmaceutical manufacturers and reallocating resources within soci-

ety. By adjusting taxes appropriately, policymakers can ensure that manufacturers—often large multinational corporations—contribute more fairly to public revenue. Considering manufacturers' strategic responses to changes in taxes and regulations is important for evaluating policy effectiveness in markets with imperfect competition.

VI Conclusions

I study the transmission of government-imposed pharmacy markups and VAT rates to retail prices in the Finnish pharmaceutical market. I use a DID strategy to demonstrate that pharmaceutical manufacturers responded to a decrease in pharmacy markups by increasing their wholesale prices. This implies that the incidence of reduced pharmaceutical VAT rates falls more on the supply side than previously thought.

I also estimate a structural model of supply and demand using data from the Finnish statin market. My results confirm that firms with greater market power respond more strongly to changes in taxes and markups. My estimates imply that statin manufacturers benefited significantly from the policy change, increasing their revenues and profits 2014–2017. The results suggest that roughly two-thirds of the changes in pharmacy revenues benefited consumers (and the reimbursement system), while the rest were captured by pharmaceutical manufacturers. In two counterfactuals, I change the VAT rate and show that the government can decrease wholesale prices and increase tax revenue by decreasing or by removing the tax subsidy altogether. Tax revenues also increase more than the expenditure from increased consumer copayments and government reimbursements. The change in VAT rates compensates for the decreases in pharmacy markups and the resulting increases in wholesale prices.

My results question the sensibility of using reduced VAT rates to subsidize demand, even if the policy goal is to support poorer consumers. First, the tax subsidy likely increases the social costs of pharmaceutical care by increasing wholesale prices. Because public reimbursement systems make aggregate demand inelastic and already insure the most vulnerable consumers against price increases, wholesale prices serve as the primary proxy for true costs of pharmaceutical care in countries that rely heavily on imported pharmaceuticals. Although effectively targeting transfers by health or other socioeconomic status is not without challenges, policymakers can fund further redistribution with the additional revenue generated by removing these tax subsidies.

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A Data Sources

Table A.1 presents the data sources used in the analysis. The data from Sweden, Denmark, and Norway are used for the construction of Hausman-like instruments.

Table A.1: Data Sources

	Years	Source
Finland	1998–2017	FIMEA
Sweden	2006Q2–2017	IQVIA
Denmark	1991–2017	DLI-MI
Norway	2000–2018	Farmastat

Notes: FIMEA = Finnish Medicines Agency; PPB = (Finnish) Pharmaceutical Pricing Board; NOMA = Norwegian Medicines Agency; TLV = (Swedish) Dental and Pharmaceutical Benefits Agency.

B Additional Tables

Table A.2: Annual Absolute Financial Changes With VAT Increases

	2014	2015	2016	2017
Panel A: Base level				
Quantities	184329.74	189995.57	190795.45	202754.37
Manufacturer Revenues	37506.66	45865.38	51858.07	49338.35
Manufacturer Profits	19392.09	20980.39	21876.84	22301.73
Expenditure	60732.58	73635.10	82852.43	79165.60
VAT Revenue	5521.14	6694.10	7532.04	7196.87
Pharmacy Profits	16095.25	19159.65	21329.38	20573.07
Prices	38.12	44.74	50.51	45.87
Panel B: 2014 reform				
Quantities	452.19	545.59	581.17	527.86
Manufacturer Revenues	1427.22	1503.81	1529.58	1603.94
Manufacturer Profits	1195.46	1232.85	1218.60	1293.02
Expenditure	-971.05	-1344.22	-1630.63	-1427.15
VAT Revenue	-88.28	-122.20	-148.24	-129.74
Pharmacy Profits	-2099.99	-2478.03	-2738.16	-2637.59
Prices	-0.88	-1.17	-1.38	-1.17
Panel C: 2014 reform + 14% VAT				
Quantities	165.15	165.52	144.98	123.26
Manufacturer Revenues	458.36	440.90	424.49	478.91
Manufacturer Profits	406.31	377.52	350.78	382.78
Expenditure	-334.28	-393.51	-419.83	-382.28
VAT Revenue	1896.19	2300.48	2591.26	2478.27
Pharmacy Profits	-2638.46	-3123.33	-2298.32	-2285.40
Prices	-0.24	-0.29	-0.32	-0.29
Panel D: 2014 reform + 24% VAT				
Quantities	-561.15	-799.70	-965.07	-898.76
Manufacturer Revenues	-1707.22	-1943.53	-2063.21	-2048.06
Manufacturer Profits	-1346.41	-1523.48	-1582.70	-1637.14
Expenditure	1227.93	1935.93	2543.51	2174.86
VAT Revenue	6471.21	7932.55	8996.21	8546.44
Pharmacy Profits	-4668.87	-5431.81	-5948.07	-5809.48
Prices	1.36	1.89	2.33	1.91

Notes: This table presents the annual absolute financial changes. All units but prices are in thousands.

Table A.3: Annual Relative Financial Changes With VAT Increases

	2014	2015	2016	2017
Panel A: 2014 reform				
Quantities	0.25	0.29	0.30	0.26
Manufacturer Revenues	3.81	3.28	2.95	3.25
Manufacturer Profits	6.16	5.88	5.57	5.80
Expenditure	-1.60	-1.83	-1.97	-1.80
VAT Revenue	-1.60	-1.83	-1.97	-1.80
Pharmacy Profits	-13.05	-12.93	-12.84	-12.82
Prices	-2.32	-2.61	-2.74	-2.56
Panel B: 2014 reform + 14% VAT				
Quantities	0.09	0.09	0.08	0.06
Manufacturer Revenues	1.22	0.96	0.82	0.97
Manufacturer Profits	2.10	1.80	1.60	1.72
Expenditure	-0.55	-0.53	-0.51	-0.48
VAT Revenue	34.34	34.37	34.40	34.44
Pharmacy Profits	-16.39	-16.30	-10.78	-11.11
Prices	-0.64	-0.66	-0.64	-0.63
Panel C: 2014 reform + 24% VAT				
Quantities	-0.30	-0.42	-0.51	-0.44
Manufacturer Revenues	-4.55	-4.24	-3.98	-4.15
Manufacturer Profits	-6.94	-7.26	-7.23	-7.34
Expenditure	2.02	2.63	3.07	2.75
VAT Revenue	117.21	118.50	119.44	118.75
Pharmacy Profits	-29.01	-28.35	-27.89	-28.24
Prices	3.56	4.23	4.62	4.17

Notes: This table presents the annual relative financial changes. All units are in percentages.

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