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ABSTRACT

Regulation of Pharmaceutical Prices: Evidence from a Reference Price Reform in Denmark*

Reference prices constitute a main determinant of patient health care reimbursement in many countries. We study the effects of a change from an "external" (based on a basket of prices in other countries) to an "internal" (based on comparable domestic products) reference price system. We find that while our estimated consumer compensating variation is small, the reform led to substantial reductions in list and reference prices as well as co-payments, and to sizeable decreases in overall producer revenues, health care expenditures, and co-payments. These effects differ markedly between branded drugs, generics, and parallel imports with health care expenditures and producer revenues decreasing and co-payments increasing most for branded drugs. The reform also induced consumers to substitute from branded drugs – for which they have strong preferences – to generics and parallel imports. This substitution also explains the small increase in consumer welfare despite a substantial decrease in expenditures.

JEL Classification: I18, C23

Keywords: pharmaceutical markets, regulation, co-payments, reference pricing, welfare effects

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1 Introduction

A steadily growing life expectancy, aging populations, and the increasing cost of medical treatments have induced policy makers to introduce various cost containment tools. Reference pricing, where patients are reimbursed a fraction of the retail price when buying a prescription drug, constitutes a particularly widely embraced approach (Berndt and Dubois 2012; Espín et al. 2011; López-Casasnovas and Puig-Junoy 2000).

While existing studies have shown that reference pricing effectively curtails prices of prescription drugs (Aronsson et al. 2001; Brekke et al. 2009, 2011; Kanavos et al. 2008; Pavcnik 2002; Puig-Junoy 2007), a hitherto empirically unanswered issue is to what extent differences in the *design* of reference pricing systems affect market outcomes. A particularly relevant question here is whether reference prices should be determined "externally", through a basket of similar products in other countries, or "internally", through prices of similar domestic products. We address that question by estimating the effects of a reference pricing reform in Denmark, a country that switched from external to internal reference pricing in April 2005. In Denmark, patients are reimbursed 80% of the reference price. The difference between the list price and the reimbursement — the co-payment — is paid by the consumers. Danish patients always co-pay and the reform did not change the 80% reimbursement rate.

Since the Danish reference pricing reform affected all drugs equally — branded drugs (on– and off–patent), generics, and parallel imports[1] — we study to what extent these different types of products were differently affected by the reference pricing reform. We confine our analysis to statins which currently constitute the best-selling drugs in terms of sales both in Denmark and worldwide. Statins treat high levels of cholesterol and are used to decrease mortality and morbidity of patients with cardiovascular diseases.

We find that the *design* of reference price systems matters substantially for prices and demand. In particular, the switch from external to internal reference pricing reduced both list prices, reference prices, and consumer co-payments by around 22%. There are substantial differences between the three types of drugs we consider: prices fall most for generics followed by parallel imports and branded drugs, where in the latter case consumer co-payments actually increased. Overall producer revenue and public expenditures both decrease by around 19% while consumer expenditures decrease by 17% as a consequence of the reform.

[1] Parallel imports are drugs that parallel importers, independent commercial agents, buy in a low-price country, re-package, re-label, and distribute in a high-price country. Parallel importing is legal in the European Union.

As the first paper to apply a structural demand estimation that is based on a consumer utility function we are able to calculate a proper measure of consumer welfare changes induced by a modification of pharmaceutical pricing regulation. We estimate an annual total consumer compensating variation (the amount government would need to pay consumers for them to accept foregoing the reform) of six million Danish krones (DKK) — around one million US dollars per year. The relatively small increase in consumer welfare seems at odds with our finding of a dramatic decrease in total patient co-payments. Using changes in co-payments as a welfare measure alone, however, ignores that the reform makes consumers more price sensitive due to increasing co-payments for the more expensive branded drugs which in turn leads them to substitute away from their otherwise preferred branded drugs. Such consumer welfare-decreasing substitution effects go unnoticed if total patient co-payments alone are used as a welfare measure as in previous studies (e.g. Brekke et al. 2011; Granlund 2010).

The paper closest to ours is Brekke et al. (2011) who exploit a quasi-experimental transition from price cap regulation to endogenous reference pricing that affected a subset of high volume off-patent drugs in Norway in 2003. They find that the switch from price cap regulation to reference pricing significantly decreased both prices for branded drugs and generics and that the change lead to reductions in the market shares of branded drugs. Brekke et al. (2011) constitutes one of few papers that study both price and demand effects of a pharmaceutical pricing reform. Their demand estimation is, however, restrictive in that it employs linear market share equations relying on the implicit assumption that all products under consideration are perfect substitutes.

We attempt to generate more flexible and hence more reliable estimates of the causal effects that the reform of Danish reference price design may have entailed on the demand for statins. The counter-factual experiment we conduct is to ask what the reform effects would have been had it occurred in the period before it was actually put in place. The advantage of this strategy is that we can effectively "filter out" factors other than the reform that may have simultaneously affected pharmaceutical market outcomes. Specifically, we first estimate a flexible logit-type demand model (Berry 1994; Berry et al. 1995) that allows for both horizontal and vertical product differentiation as well as for arbitrary substitution patterns between products by allowing for consumer-specific heterogeneity in drug demand. Second, we estimate pricing equations to predict the counter-factual prices of drugs had the reform taken place before it actually

did. Finally, we use our estimated pricing and structural demand parameters to compute counter-factual demand which allows us to calculate total changes in demand, consumer expenditures, producer revenues as well as consumer welfare.

The paper proceeds as follows: Section 2 describes the Danish pharmaceutical market and the institutional settings of the reference price reform, Section 3 describes our data set, Section 4 describes the empirical strategy, Section 5 summarizes our estimation results, and Section 6 concludes.

2 The Danish market for pharmaceutical products

As in other European countries, the market for pharmaceutical products in Denmark is regulated. Denmark follows EU regulations regarding product authorization. Product pricing, reimbursement rules, and the regulation of pharmacies are national matters.

The pricing of pharmaceutical products in Denmark is free.[2] Changes in pharmacy purchase prices are notified to and evaluated by the Danish Medicines Agency (DKMA). The agency updates prices every 14 days and makes them publicly available online. Prices are identical nationwide.

In Denmark, pharmacists must first offer the patient the cheapest product within a group of substitutes unless the prescription explicitly requires no substitution, which is the case for just five percent of all prescriptions. The patient may then decide herself whether or not she buys the cheapest product or a substitute at a higher price and a higher co-payment. Other relevant market features are that (i) Denmark maintains a universal health care system that is financed through general tax revenues, (ii) that advertising prescription drugs to patients is prohibited and (iii) that detailing is regulated. Detailing it is mainly used for new products and not for established drugs, such as the ones in our analysis.

The reform that this paper investigates involves the change in the way reference prices are calculated. On April 1, 2005, Denmark changed from external to internal reference pricing for all prescription-only pharmaceuticals independent of their patent status. The classification of products into substitution groups remained unchanged. In Denmark, patients may only substitute among products with the same active substance, administration form, strength, and similar package size, where package size may not vary by

[2]There exists one fairly loose restriction, however, by that drugs for which an analogous product exists cannot be reimbursed if its price is more than 20 per cent higher than the price of the analogous drug.

more than ten percent within substitution groups.

Before the reform reference prices were based on average prices in the EU-15 member states, excluding Greece, Luxembourg, Spain, and Portugal. The reference price for a given product was the lowest cross-state average price among products belonging to the same substitution group. However, if a product's list price was below the EU average, the reference price was set equal to the list price. After the reform, the reference price was set equal to the lowest domestic list price out of all products belonging to the same substitution group.

Around the time of the reform there were other events happening that might have influenced the behavior of the market participants. We grouped these events and divided our observed data into six different periods, which are summarized in Appendix A. Our main relevant dates were set by the Danish government. In May 2004, the Danish parliament ratified the new reimbursement law making it public in June 2004. On April 1, 2005, the law was implemented. However, it is likely that information regarding changes in reimbursement rules had been at the disposal of market participants prior to these two legislatively determined dates. On September 17, 2003, the Danish Minister of Health announced the assembly of a group of experts with the aim of changing the existing reimbursement system to strengthen competition. Moreover, as a member of the working group, the Danish Association of the Pharmaceutical Industry (Lægemiddel Industri Foreningen, LIF) launched the idea of changing the way reference prices are calculated, as was eventually adopted in April 2005. Between May 2001 and April 2003, LIF maintained a voluntarily agreement on price ceilings. However, not all members complied with the agreement. After its expiration in 2003, LIF announced a continuation of the price ceiling for another two years. This was a unilateral announcement on the side of LIF rather than an official agreement with the Danish Ministry of Health.[3] Finally, the Danish Ministry of Health and LIF again signed an agreement on a price ceiling in October 2006.

[3]Notwithstanding, we cannot exclude the possibility that the LIF announcement allowed producers in the market to coordinate on higher prices levels (Knittel and Stango, 2003). However, uncertainty regarding the credibility of the LIF announcement, as well as the volatile market structure following the patent expiration of a popular product, Zocor, in 2001, suggest that price coordination was difficult to sustain. For this reason, we interpret the price development as being the result of the announced reform, but we are not able to separate the effects of the reform from the possible effects of the LIF announcement.

Our analysis focuses on the base period (May 03, 2001 until April 14, 2003) and the implementation period (April 01, 2005 to September 25, 2006). Our base period is the time between the working group assembly and the ratification in parliament. It serves as a base because no reliable information about prospective changes in the reimbursement system was publicly or privately available and because the number of firms as well as prices remained stable. Our treatment period covers the actual implementation of the reform. We discard the two LIF agreement periods as well as the adjustment period after the expiration of the first LIF agreement to avoid including effects other than the actual reform. We also discard the announcement period because firms were informed about the new legislation which allowed them to prepare for a new competitive setting.

3 Data

Our data set contains fortnightly prices and sales of statins for the period between February 2003 and June 2006. We downloaded the publicly available price data from http://medicinpriser.dk/. The sales data are proprietary and were made available to us by LIF. They come with the same periodicity as the price data.

The site http://medicinpriser.dk/ contains a list of all authorized pharmaceutical products marketed in Denmark. Prices are updated every second Monday based on changes reported by producers during the last two weeks. The data base is used by general practitioners when issuing prescriptions, by hospitals for their electronic patient records, and by pharmacies to ensure nationally uniform prices for prescription drugs.

A pharmaceutical product is characterized by its name, package size, form of administration, strength, 5-level anatomical therapeutic chemical classification code (ATC code), and producer name. The ATC-code is a combination of letters and digits that precisely describes a product's active substance.

Appendix B contains a characterization of statins in terms of their ATC code. Statins are divided into eight different ATC classes, of which six are marketed in Denmark. Three of them (Simvastatin, Lovastatin, and Pravastatin) lost patent protection before our data set starts which induced generic entry to the market. Fluvastatin lost patent protection by the end of 2003 and the remaining two molecules, Atorvastatin and Rosuvastatin, are on-patent during the whole period we analyze. The post-reform reference price for these

two on-patent drugs is then determined by parallel imports.[4]

Medical practitioners in Denmark tend to regard all statins as close substitutes, at least with respect to their effects on cholesterol levels and slightly less so with respect to their resorption. When treating a patient, they follow the recommendations issued by the Institut for Rational Farmakoterapi (IRF, an institution under the Danish Medicines Agency that seeks to promote the most rational use of medical products) and simultaneously choose the active ingredient and dosage. It is not clear a priori if and to what extent Danish medical doctors and patients are price sensitive. IRF does, however, issue recommendations to substitute one product by another if (i) it has been demonstrated in clinical studies that the effects are identical, and (ii) one of the products is substantially cheaper than the other.

Table 3 presents a descriptive overview of prices and sales of statins. To make the different strengths, package sizes, and active ingredients comparable we converted prices and quantities into Defined Daily Dosages (DDD).[5]

Prices are in Danish crowns (DKK) and are deflated using the consumer price index with the year 2005 as the basis. The average list price of statins is 7.8 DKK per DDD across all periods and products. Average reference prices are 6.1 DKK and consumer co-payments are 2.9 DKK. These prices differ substantially across the three different types of drugs. Branded drugs are most expensive with an average list price of 12.2 DKK. Generics are cheapest and cost on average 3.6 DKK while parallel imported drugs cost on average 11 DKK.

All prices decreased from the base to the implementation period on average. This decrease was stronger for list prices than for co-payments. The decline in list prices from the base to the implementation period is smaller for branded drugs than for generics or parallel imports. Co-payments even increased for branded drugs, on average from 4.6 to 5.8 DKK per DDD.

Sales are on average highest for generics, followed by brands and parallel imports. From the base to implementation period, sales for generics and parallel imports increased on average and decreased for

[4]Although parallel importing of generics is possible, most parallel imports in our data are branded products.

[5]We cannot exclude that our DDD normalization suppresses potential non-linearities in pricing. It is unclear,, however, how that would affect our analysis. In addition, such a problem would only materialize if pricing strategies changed with the reform, an issue for which we do not find any evidence. Moreover, any time-invariant differences in pricing strategies will be accounted for by our product fixed effects.

Table 1: Prices and sales for statins

	All			Brand			Generics			Parallel Imports		
	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.
All periods												
List price	7.79	5.18	6.78	12.15	11.28	5.98	3.61	2.52	3.28	11.00	9.76	7.81
Reference price	6.09	3.43	6.16	8.91	7.38	6.20	2.91	1.90	3.00	9.11	6.90	7.51
Consumer co-payment	2.92	1.47	4.17	5.02	2.78	5.37	1.28	0.72	1.70	3.71	2.67	4.68
Quantities (in 1'000 DDD)	24.44	2.19	114.30	20.06	5.73	36.46	34.48	1.44	157.31	9.53	0.63	55.59
Obs.	13'861			3'907			6'633			3'321		
Base period												
List price	9.37	7.57	6.51	12.56	11.57	5.38	5.10	3.64	4.06	12.69	13.05	6.87
Reference price	7.19	4.33	5.93	9.95	9.21	5.46	3.84	2.45	3.59	9.51	10.07	6.70
Consumer co-payment	3.62	2.09	4.82	4.60	2.75	5.38	2.02	1.15	2.61	5.08	3.40	6.00
Quantities (in 1'000 DDD)	18.77	2.41	61.01	18.98	6.47	31.40	25.09	1.36	85.80	8.57	0.74	27.67
Obs.	2'524			744			1'090			690		
Implementation period												
List price	5.81	3.41	6.11	11.57	9.57	6.23	2.58	1.95	2.25	7.32	4.09	7.23
Reference price	4.21	2.03	5.39	7.26	6.75	6.58	2.01	1.38	2.07	6.63	3.73	7.19
Consumer co-payment	2.44	1.06	3.67	5.75	2.92	5.42	0.97	0.60	1.02	2.02	1.30	2.14
Quantities (in 1'000 DDD)	26.54	1.60	125.80	18.18	4.83	34.17	35.72	1.05	164.81	9.52	0.44	32.38
Obs.	4'963			1'340			2'781			842		

Prices per Defined Daily Dose (DDD) in June 2005 DKK, where 1 DKK = 0.165 US dollar. Quantities measure as sold DDD.

brands.

Appendix C summarizes other market and product characteristics such as the number of products on the market, the number of firms active in the market, average package size, and average strength. It shows that half of the products are generics and that there are more producers of generics than brand manufacturers or parallel importers. We observe an increase in the number of generic products from the base to implementation period (from 54.5 to 70.3 on average) and a decrease in the number of branded and parallel imported products. The products we consider are all pills, coated pills or capsules. The median package size is 98 pills, and the median strength is 20 milligram of active substance per pill. These characteristics do not vary much between the base and the implementation period.

4 Empirical strategy

Our empirical strategy to identify the effects of the reimbursement reform on prices and demand proceeds in three steps. We first estimate a structural demand model that maps observed and unobserved product and consumer characteristics to product sales. Second, we estimate a reduced-form pricing equation that studies to what extent prices changed due to the reform. This estimation generates the prices that would have been observed had the reform taken place in the base period already. Third, we use our estimated counter-factual prices and plug them into our demand model for the base period, the period before the reform. This generates counter-factual demand for the base period given our predicted counter-factual prices for the base period. The reform effects are identified by comparing these counter-factuals with observed base period market outcomes.

4.1 Demand Model

LMAs, as many other drugs, are both vertically and horizontally differentiated products. In our model, we account for vertical differentiation by including product brand names and package size as observable characteristics. An idiosyncratic error term allows for horizontal differentiation.

To estimate the demand for statins we employ a random coefficients logit model due to Berry et al. (1995). This model assumes that in every time period each individual consumer i chooses product j

that maximizes her utility.[6] Omitting the time index t for notational convenience, her utility function is:

$$U_{ij} = \delta_j + \sigma_p p_j^c \nu_{ij} + \varepsilon_{ij}, \tag{1}$$

where all consumers obtain mean utility δ_j, which is common to all consumers and individual-specific utility $\sigma_p p_j^c \nu_i + \varepsilon_{ij}$. The term p_j^c denotes patient co-payment. Importantly, we allow for variation of consumer preferences for price in the population by including the term $\sigma_p p_j^c \nu_i$. Own-price and cross-price elasticities may hence vary across individuals which generates much more plausible price elasticity estimates compared to the computationally less burdensome simple logit and nested logit models for differentiated products demand (Berry et al. 1995). To identify consumer preferences regarding price, we assume ν_{ij} to be drawn from a standard normal distribution with standard deviation σ_p, a parameter that is to be estimated. If σ_p is insignificantly different from zero, the model collapses into the simple logit model (Berry 1994). The idiosyncratic random error term ε_{ij} is assumed to be i.i.d. Gumbel distributed.

We decompose mean utility into

$$\delta_j = \boldsymbol{x_j}\boldsymbol{\beta} - \alpha p_j^c + \xi_j, \tag{2}$$

where $\boldsymbol{x_j}$ denotes a vector of observed product characteristics and ξ_j is an unobservable product characteristic.[7]

Vector $\boldsymbol{x_j}$ includes sets of dummy variables for product names, strength of the active ingredient, and package size. These three characteristics implicitly define substitution groups which are set by the regulator and hence impose a soft restriction on the choice set. We further include monthly dummy variables to control for seasonal variation as discussed by Ockene et al. (2004) and Tung et al. (2009) as well as time period dummies.

The assumption that consumers are utility-maximizers combined with the assumption that ε_{ij} is i.i.d. Gumbel distributed leads to the following market share equation (Berry 1994):

$$s_j(\boldsymbol{x_i}; \boldsymbol{\theta}) = \int_\nu \frac{exp(\delta_j + \sigma_p p_j^c \nu_i)}{1 + \sum_J exp(\delta_j + \sigma_p p_j^c \nu_i)} dF_\nu(\nu), \tag{3}$$

[6] Due to the aggregate nature of our product-level data, we assume that the consumer entity is a joint physician-patient unit and, consequently, abstract from possible agency problems. The assumption holds if physicians act in the interest of their individual patients. See Dunn (2012) for a recent example using a similar assumption in modeling demand for statins.

[7] Note that the mapping between the co-payment and consumer utility follows from the combination of the mean, $p_j^c \alpha$, and individual-specific utility terms, $p_j^c \sigma_p \nu_{ij}$: $p_j^c(\sigma_p \nu_{ij} - \alpha)$.

where vector $\boldsymbol{\theta}$ contains the coefficient vector $\boldsymbol{\beta}$, identical for all individuals, and parameter σ_p.

To close the model, we need to define potential market size and implicitly the share of outside good $j = 0$. Consumption of the outside good provides consumers with a mean utility that we normalize to 0 ($\delta_0 = 0$). In our setting, the composite outside good consists of products that are not statins and that may reduce cholesterol level including, for example, non-statin LMAs, homeopathic products, a bicycle, or a pair of running shoes.

The price of our outside good is not set in response to the prices of the inside goods, the statins. We define total market size as the amount of DDDs sold if all *potential* patients had received statins as medication. We infer the number of potential patients based on a claim by the Danish Association of Heart Patients (Hjerteforeningen, 2007) that 60% of all Danish residents between ages of 40 and 80 years have an elevated cholesterol level. At a total Danish population of 5.5 million this fraction matches well with IRF's (IRF, 2006) estimate that 2.1 million Danish residents above the age of 35 have a total cholesterol level of more than 5 mmol/l, the critical threshold above which treatment with statins is started. As we base our estimates on DDD, a daily per-patient unit, the potential market size can be computed simply as 60% of all Danish residents between the ages of 40 and 80. We employ this broad market definition to provide conservative demand estimates. Decreasing potential market size, for example, by assuming a lower fraction of people with elevated cholesterol levels, increases absolute elasticities of substitution.

The term ξ_j is unobserved by the econometrician but observed by both consumers and producers. In our setting, we think of this characteristic as quality perception in the market which might deviate from the time-invariant mean product name effect we explicitly control for. This quality perception may vary over time and can be influenced by changes in consumer information through channels such as producer publicity, post-entry clinical testing, and population product experience. Profit-maximizing producers will adjust prices to changes in ξ_j which leads to omitted variable bias in the estimated price coefficient. We address the resulting endogeneity problem by employing a set of instruments and estimating the model using GMM. Following Dubé et al. (2012) we write the objective function as a constrained optimization

problem for numerical robustness:

$$\begin{aligned} \min_{\boldsymbol{\theta}, \boldsymbol{x_j}} \quad & \boldsymbol{x_j}' \boldsymbol{Z} \boldsymbol{W} \boldsymbol{Z}' \boldsymbol{x_j} \\ & \text{subject to } s(\boldsymbol{x_j}; \boldsymbol{\theta}) = \boldsymbol{S}, \end{aligned} \tag{4}$$

where the vector $\boldsymbol{Z}$ denotes a set of optimal instrumental variables, the vector $\boldsymbol{W}$ denotes a weighting matrix, and $\boldsymbol{S}$ are the observed market shares. In the construction of the vector of optimal instruments, which closely follows Reynaert and Verboven (2012), we rely on identification arguments in Berry et al. (1995) who include variables containing information about the competitive environment. These covariates, termed "BLP instruments" hereafter, are the sums of other firms' products' characteristics (package size and strength of active ingredient) as well as the number of competitors in the market and in the relevant substitution groups. A detailed description of the identification and estimation of our model is relegated to Appendix D.

With a fully specified demand model and counter-factual prices at hand we can compute a simple monetary measure of reform effects on consumer utility, the Hicksian compensating variation. Formally, we obtain consumer compensation variation measure by solving the integral over the differences in maximum expected utilities via numerical simulation (see Small and Rosen, 1981):

$$CV = \int \frac{1}{\alpha + \nu_i} \left\{ \ln \sum_j \exp\left(\delta_j^{pre} + \sigma_p p_j^{c,pre} \nu_i\right) - \ln \sum_j \exp\left(\delta_j^{post} + \sigma_p p_j^{c,post} \nu_i\right) \right\} f\left(\boldsymbol{\nu} \mid \theta^{pre}\right) d(\boldsymbol{\nu}) \tag{5}$$

We will use the parameters of our demand model to predict counter-factual demand (superscript "post") for statins based on the counter-factual prices whose estimation we discuss in the subsequent paragraph.

4.2 Reduced-form Price Equation

The idea behind our pricing regression is to infer price changes due to the reform by regressing actual list prices on a large set of control variables, fixed effects, and a set of dummy variables for the reform. This allows us to calculate the prices that would have been observed had the reform happened in the base period.

We could in principle also compare prices in the base and the reform period to infer what products would have cost in the absence of the reform. That would, however, imply to discard products that

were unavailable either in the base or in the reform period. It would also imply forgoing to control for confounding factors such as the competitive environment in the counter-factual reform, the base period.

To identify our pricing equation we exploit the panel structure of our data. In particular, we rely on within-variation for identification by using time-invariant product name fixed effects. These fixed effects also capture important product-level market characteristics such as the time a product has been on the market. In addition, we control for seasonal within-year trends using month fixed effects and for time-invariant cost of active ingredient strength and package size by including substitution fixed effects as well as pulp and paper prices. Pulp and paper prices are input prices and affect prices but not the unobserved quality ξ_j. In addition, we include the set of BLP instruments discussed in Subsection 4.1.

We interact the reform dummy with dummy variables for the type of product, namely if it is a branded, generic, or parallel imported drug. While Pavcnik (2002), Granlund (2010), and Brekke et al. (2009, 2011) find strongest price decreases for branded products, some earlier studies provide evidence for non-decreasing prices for branded products that goes along with increased competition (Frank and Salkever, 1997, Grabowski and Vernon, 1992, Regan, 2008). This has been labeled the "generic competition paradox" (Scherer, 1993). The intuition here is that brands may themselves differentiate even further and only target low-elasticity consumers to avoid facing tougher competition.

While many studies explore the link between reference pricing and competition between brands and generics, we are able to identify and differentiate a third group, parallel imports. The efficacy of parallel importing as a tool to improve price competition has been a highly debated topic.

We consider list prices as the relevant price outcome as these are the prices producers set. They mechanically define co-payments and reference prices after the reform. We use the linear panel specification

$$\ln p_{jt} = \gamma_1 D_t + \gamma_2 D_t * \mu_b + \gamma_3 D_t * \mu_{PI} + \mu_j + \mu_m + \mu_s + \gamma_4 N_t + \gamma_5 N_{st} + \boldsymbol{z_{jt}\gamma} + \varepsilon_{jt}, \tag{6}$$

where the dependent variable is the log list price per DDD of product j at time t. D_t equals one in the implementation period and zero in the base period. Further indicator variables are denoted by μ, where subscript b indexes brands, PI parallel imports, j products, m months, and s substitution groups. The specification also controls for the number of products in the market, N_t, the number of products in product j's substitution group, N_{st}, the set of BLP instruments, and production cost factors. The latter variables

are stacked in vector $\boldsymbol{z_{jt}}$. The term ε_{jt} denotes an idiosyncratic shock.

From our estimation of Equation (6) we calculate counter-factual product prices in the base period, period BP:

$$\hat{p}_{jBP} = exp(\hat{\gamma}_1 + \hat{\gamma}_2 * \hat{\mu}_b + \hat{\gamma}_3 * \hat{\mu}_{PI} + \hat{\mu}_j + \hat{\mu}_m + \hat{\mu}_s + \hat{\gamma}_4 N_{BP} + \hat{\gamma}_5 N_{sBP} + \boldsymbol{z_{jBP}}\hat{\boldsymbol{\gamma}}). \tag{7}$$

5 Estimation results

Our estimation results fall in three parts. We first discuss our demand model, proceed with our price estimations, and finally evaluate the reform effects on prices, demand, and consumer surplus.

5.1 Demand Parameters

Table 2 reports the estimated coefficients and the implied price elasticities with respect to consumer co-payment for three alternative specifications of our demand model. The left columns present OLS logit results where we assume that consumers have homogeneous preferences with respect to patient co-payment ($\sigma_p = 0$) and that prices are exogenous to demand. The middle columns show IV logit results where we instrument prices. The right column displays random coefficients logit model results, our main and preferred specification.

We estimate a negative and significant co-payment coefficient in the OLS Logit model and a mean own-price elasticity of -.22. We refer to Berry (1994) for a derivation of the price elasticities for our three models. While low in absolute terms, this simple model obtains the correct negative sign for the co-payment coefficient. Once we instrument prices, identification of the co-payment improves, as measured in terms of t-values, substantially and the coefficient more than doubles as opposed to the non-instrumented estimates. The corresponding mean own-price elasticity is -.59. We report our first stage regression results, our regression of our endogenous variable list price on the instruments and the exogenous variables, in Appendix E. The tests for joint instrument significance are all substantially above the critical value of ten that Stock et al. (2002) suggest.

As we have little reason to believe that all individuals in Denmark are equally sensitive to price changes, we drop the assumption that $\sigma_p = 0$ in the RC Logit model. This full model reveals significantly more

Table 2: Logit and random coefficient logit demand

	OLS Logit	IV Logit	RC Logit - MPEC	
			Mean	Std. dev.
Co-payment	-.14*** (.006)	-.39*** (.033)	-1.54*** (.271)	.54*** (.086)
Package Size	.02*** (.001)	.02*** (.0006)	.02*** (.001)	
Strength	.01*** (.001)	-.003 (.003)	-.004 (.003)	
Constant	-9.85*** (.133)	-6.52*** (.471)	-6.14*** (.594)	
R^2	.42	.35		
# obs.	13'861	13'861	13'861	
η_j (mean)	-.22	-.59	-1.19	

Notes: Robust standard errors in parentheses. Product name, month, and time period dummies are included. F-value in the first-stage regression of IV Logit: 136.16. 5'000 modified latin hypercube sampling draws used to simulate market shares in the random coefficients logit model. Elasticities η_j are market share weighted mean elasticities in the base period.

price elastic demand with a mean estimate of -1.54 and a corresponding standard deviation of .54. The implied mean elasticities are double the IV Logit ones.

Our specification does not include a single dummy for branded drugs. We include a set of 42 product name dummies instead, the corresponding coefficient estimates are, however, not displayed for brevity. Averaging over these brand name dummies for branded drugs, parallel imports, and generics shows that the coefficients related to branded drugs are four times larger than for the other two drugs types, whose coefficients on name dummies are fairly similar.

Our estimates suggest that consumers are more price elastic than what is found in most of the existing literature on pharmaceuticals demand (Gemmill et al., 2007, for a survey). This is not surprising as an external reference price mechanism was in place in Denmark before the reform. Even with external reference prices, consumers were faced with the choice between buying either cheaper generics and parallel imported drugs or the more expensive branded drugs and, hence, they were more price-sensitive than in markets with little co-payment.

In all three models, the coefficient estimates on month indicator variables are in line with first evidence by Ockene et al. (2004) and Tung et al. (2009), who find that lipid levels are low in the summer and high

in the winter. The corresponding coefficient estimates are not displayed in the table for brevity.

5.2 Prices

Our next step is to calculate the change in prices the reform induced. We run a total of six alternative pricing regressions. Table 3 presents the coefficient estimates in the order of increasing numbers of control variables. We shall use the full specification, depicted in column (6), to compute all reform effects in the following subsections. The estimation sample contains observations on all products on the market in the base period and the implementation period. To take into account potential serial correlation we compute standard errors that are robust to autocorrelation and heteroskedasticity and clustered at the product level.

In all specifications, we obtain a negative average effect of the reform on list prices. Specification (6) implies that, on average, the reform decreases generic prices by 35.8%, brand prices by 7.3%, and parallel import prices by 18.7%.[8] Specification (5) excludes only the brand and parallel import interaction terms. It estimates the reform effect on list prices over all types of products at -21.4%.

Specification (1) comes with a minimum of control variables and significantly overestimates the average reform effect with -40.6%. Adding BLP instruments as basic controls for the competitive environment in specification (2) reduces the bias to some extent. Both coefficients of the numbers of products in the market and in substitution groups as a further competition control variable in specification (3) are significant and negative, as expected based on standard oligopoly theory. Here, the estimated reform effect doubles to -78.8% which is a sign for substitution group specific effects of the reform on product entry and exit, i.e. selective entry and exit. Including these continuous control variables, however, we cannot discriminate between changes and time-invariant levels in the numbers of products in substitution groups. Therefore, we include further substitution group fixed effects in specification (4). The latter almost nullify the estimates of γ_4 and γ_5 in specification (3) and the bias of the reform effect estimate is further reduced. Finally, in specifications (5) and (6), we add an input cost index (pulp and paper) and product name fixed effects to control for time-invariant levels of product quality. The latter should alleviate concerns that selection may

[8] We use a log-linear specification with dummy explanatory variable D_t and so the percentage effect of the reform on list prices is defined as $\exp\left(\hat{\gamma}_1 - \frac{1}{2}V(\hat{\gamma}_1) - 1\right) \times 100$ (see Kennedy, 1981).

Table 3: Price regressions

	$\ln$ *retail price* ($N = 7487$)					
	(1)	(2)	(3)	(4)	(5)	(6)
Reform (γ_1)	-.52*** (.057)	-.48*** (.111)	-1.53*** (.194)	-.46*** (.061)	-.24*** (.046)	-.44*** (.072)
Reform × Brand (γ_2)						.37*** (.077)
Reform × Parallel Import (γ_3)						.24** (.094)
No. Products in Market (γ_4)			-.10*** (.023)	-.001 (.004)	-.002 (.004)	-.001 (.004)
No. Products in Subst. group (γ_5)			-.12*** (.012)	-.07*** (.015)	-0.007 (.013)	-0.02* (.012)
Pulp & Paper × Product name	No	No	No	No	Yes	Yes
BLP Instruments	No	Yes	Yes	Yes	Yes	Yes
Fixed effects						
Product name	No	No	No	No	Yes	Yes
Substitution group	No	No	No	Yes	Yes	Yes
Month	Yes	Yes	Yes	Yes	Yes	Yes
Constant	2.10*** (.064)	2.02*** (.402)	4.35*** (.526)	1.64*** (.254)	-4.798*** (1.592)	-5.25** (1.728)
R^2	.10	.13	.35	.57	.89	.89

Notes: The table reports linear dummy variable regression estimates of the coefficients in Equation (1). Values between parentheses are robust standard errors clustered at the product level. The estimation sample contains only the base and the implementation period.
*** p< 0.01, ** p< 0.05, * p < 0.1

confound our estimates of the reform effect.

Our results can be explained by the mechanisms similar to the ones suggested by Brekke et al. (2011). The Danish reform strengthened firms incentives to decrease prices by giving price setters the possibility to influence the market reference price. As the Danish reform entailed a change within an existing reference price system, the size of its impact on list prices and consumer co-payments had been an open empirical question. Our results provide a first attempt to quantify these effects.

5.3 Reform Effects on Prices, Demand, and Consumer Surplus

Our demand estimates and our estimates for counter-factual prices form the backbone of our calculation of counter-factual demand and consumer surplus. The flexibility of our demand model allows us to take into account consumers' substitution behavior caused by our estimated list price changes from which we infer the induced reference price changes and patient co-payments.

Recall that, after the reform, the reference price is defined by the lowest price in a given substitution group. A strong price decrease for low-price generics paired with a weaker price decrease for high-price brands will lead to an increase in consumer co-payment for brands. Hence, we expect the reform to be highly effective in pushing consumers to substitute away from brands towards generics and parallel imports.

Table 4 reports absolute and percentage differences between our observed market outcomes in the base period and our predicted counter-factual market outcomes had the reform already been implemented in the base period. It shows that overall list prices decrease by 21.9%, where the largest decrease is accounted for by generics with 46.4%. List prices for parallel imports decrease significantly less with 22.1% but, most remarkably, brand list prices decrease only by 7.2%. While the latter finding falls short of the generic competition paradox whereby increased competition causes increasing prices of branded drugs (Frank and Salkever, 1997; Grabowski and Vernon, 1992; Regan, 2008), our results run counter to the findings of Pavcnik (2002) and Brekke et al. (2009, 2011) who find a stronger decrease in list prices for branded products. We should keep in mind, however, that the Danish reform has not been a full switch to reference pricing but only a change in the design of an existing reference price system.

Co-payments decrease significantly both for generics and parallel imports while they increase for branded products. As the final purchase decision is with the consumer facing co-payments, these predicted effects

Table 4: Reform effects on market shares, expenditures, and consumer welfare

	All		Generics		Parallel Imports		Brands	
	Δ	Δ%	Δ	Δ%	Δ	Δ%	Δ	Δ%
In DKK (2005) per DDD								
List price	-2.05	-21.91	-2.36	-46.37	-2.80	-22.10	-.90	-7.20
Reference price	-1.61	-22.36	-1.54	-40.33	-1.80	-18.94	-1.52	-15.23
Consumer co-payment	-.77	-21.21	-1.12	-55.53	-1.36	-26.84	.31	6.71
In 1'000 DDD per year								
Quantities	5'819	9.43	9'212	26.98	1'299	28.37	-4'691	-26.08
In 1'000 DKK (2005) per year								
Producer revenue	-67'834	-18.83	-7'277	-8.97	7'282	14.07	-67'839	-29.23
Government expenditures	-52'556	-19.48	-10'766	-20.24	4'641	11.76	-46'431	-27.17
Consumer expenditures	-15'277	-16.99	3'488	24.83	2'641	22.42	-21'407	-34.91
Consumer surplus	6'142.57							

Price changes are computed from base to implementation period, based on predicted prices from specification (6) in Table 5. All other changes from base period to counterfactual implementation in base period, using estimated parameters of random coefficient logit to predict counterfactual market shares and consumer surplus. All figures in June 2005 DKK, where 1 DKK = .165 US dollars.

should induce a significant shift in demand away from branded drugs. This mechanism helps reducing expenditures even if brand list prices do not decrease significantly after the reform. Consumers substitute towards generics that witnessed large price decreases from consistently more expensive branded products. This asymmetric change in co-payments is due to the asymmetric changes in list prices but comparably uniform changes in reference prices across drug types. The quite uniform reference price changes are due to the fact that, in the market for statins, most substitution groups include both branded and generic products.

Indeed, we find that the demand for generics increases by 30% and for parallel imports by 28.4%. Demand for branded products decreases by 26.1%. These results are in line with Brekke et al. (2011) and demonstrate the power of a market-based competition-strengthening mechanism in inducing consumer switching to cheaper products.

Overall government and consumer expenditures decrease by 18.8% and 17%, respectively. Producers obtain 19.5% less revenue. The largest loss in revenues of 29.2% is incurred by branded producers. As the revenues for parallel imports increase by 14.1%, the reform has had a beneficial impact for parallel importers.

The significant decrease in consumer expenditures can be explained mostly by consumers' switching from

brands to generics and parallel imports. In addition, the population of potential consumers experienced a utility gain of 6'142'571 DKK (1'013'524 USD) per year. Given that the reform entailed a substantial total co-payment decrease for generic drugs and parallel imports, this may seem surprising. However, the reform led to co-payment increases for branded drugs which forced consumers to substitute away from branded drugs, for which they have strong preferences as indicated by the large coefficients on the name dummies for branded drugs, towards generics and parallel imports. Our finding of relatively small changes in consumer welfare shows that using consumer expenditures, as in Brekke et al. (2011) or Granlund (2010), as a proxy for patient welfare may lead to an overestimation of the reform effects. Patient expenditures do, however, not account for welfare losses due to substitution away from an otherwise preferred product.

We hence find that the Danish pricing reform has been largely successful in decreasing public expenditures and consumer expenditures. It also incentivized consumers to switch to generics or parallel imports. While significant utility gains are realized on the national level, patient welfare gains are low. Producers of branded drugs incur substantial revenue losses. We approximate the associated loss in *total* welfare by the sum of the compensating variation and changes in producer revenue as 56'473'530 DKK (9'318'132 USD).

6 Conclusions

Reference pricing constitutes a widely adopted cost containment tool used by governments to curb expenditures for pharmaceuticals. While it is well documented that reference pricing drives down pharmaceutical prices, little is known about the design of such systems.

This paper demonstrated that the design of reference price systems may substantially impact market outcomes. It analyzed the extent to which a switch from external reference pricing, where reference prices are determined based on prices of similar products in other countries, to internal reference pricing, where the price of the cheapest domestic substitute constitutes the reference price, matters for prices and demand.

We used product-level data to study the effects of a reference pricing reform in Denmark in April 2005 when the country substituted external for internal reference pricing. This reform affected all prescription drugs independent of patent status. The focus of our analysis were statins which constitute blockbuster drugs both in Denmark and worldwide.

Our analysis showed that list prices, reference prices, and consumer co-payments all decrease by around 22% due to the switch to internal reference pricing. These changes are quite unevenly distributed across different types of drugs. Prices decreased most substantially for generics where consumer co-payments declined by as much as 56%. Prices for parallel imported drugs also decreased significantly while prices for branded drugs changed comparatively little. Consumer co-payments for branded drugs increased by seven percent — a result which can be explained by a more pronounced decrease in reference prices relative to list prices for this type of drugs.

We used these predicted reform-induced changes in prices to analyze changes in drug demand caused by the reform. To this end we derived a structural model of the demand for statins that allowed us to predict counter-factual drug demand and to calculate consumer welfare effects due to the reform.

Our estimates indicate an overall increase in statins demand associated with the reform. These demand changes were again unevenly distributed across alternative drug types. Demand increased most for generic drugs, by 27%. Parallel imported drugs encountered a similarly large increase while the demand for branded drugs decreased by 26%. The switch to internal reference pricing hence induced patients to substitute towards generic and parallel imported drugs.

Combining price and demand effects we estimated an overall average decrease in producer revenue by 19%, a decrease in health care expenditures by 19%, and a decrease in consumer expenditures by 17%. Parallel importers benefited most from the reform. Their overall revenues increased by 14% which reflects the relatively small decrease in prices combined with a relatively large increase in demand. Generic producers revenues decrease slightly by nine percent while those of branded drugs decreased by as much as 29%.

We also found that health care expenditures decreased by 20% for generics and by 27% for branded drugs while they increased by 12% for parallel imported products. These results indicate that the reduction in reference prices, which constitutes a key determinant of pharmaceutical cost reimbursement, compensates the associated increase in demand. We come to an opposite conclusion for parallel imported drugs while the reduction in health expenditures for branded drugs follows directly from falling prices and demand.

Consumer expenditures also decreased as a consequence of the reform, by 17% on average. This reduction is primarily driven by the massive decrease in consumer expenditures for branded drugs (35%).

Consumer expenditures increased, however, by a quarter for generic drugs and by 22% for parallel imported drugs. In both cases, an increase in demand for the respective type of drugs over-compensates the reform-associated reductions in reference prices.

Our structural estimation of drug demand allows us to calculate consumers' compensating variation, our measure of consumer welfare. It represents the amount patients would need to be compensated for to maintain their level of utility after foregoing the reform. Our estimate for the compensating variation is six million DKK (one million USD) per year. This may seem small given the comparatively large reductions in consumer expenditures, but approximating consumer welfare by patient expenditures ignores welfare losses induced by substitution from a preferred product (branded drugs) to cheaper alternatives (generics and parallel imports).

The key result of our analysis is that not only the introduction of reference pricing as such — as shown by previous empirical studies such as Brekke et al. (2011) — may have dramatic consequences for market outcomes but that the *design* of reference pricing systems may also have substantial impacts on producers, patients, and government health care expenditures. In particular, our paper shows that a switch from external to internal reference pricing may effectively stimulate substitution away from branded drugs and reduce health care expenditures. It may, however, not lead to a substantial increase in consumer surplus.

While the present paper confined itself to the analysis of a chronic disease, future research will extent the analysis to an acute treatment like an infectious disease. We speculate that the reform effects may be considerably smaller for an acute treatment since patients may be substantially less price elastic.

Furthermore, adverse regulatory impacts on producers' static profits may lead to dynamic firm reactions, for example a reduction of research and development expenditures. Innovation is an important driver of consumer welfare in pharmaceutical markets. While beyond the scope of this paper, investigating the trade-off between static and dynamic objectives in regulatory policies for research-intensive industries is an important research agenda.

Appendix A: Summary of events related to changes in the Danish reimbursement system

LIF Agreement	May 03 2001 Apr. 14 2003	Since 2001 LIF members and the Danish Ministry of Health have an agreement on price ceiling running until 2005. Not all LIF members comply with the agreement.
Adjustment	Apr. 28 2003 Sep. 01 2003	The Danish Medicine Agency starts updating pharmaceutical prices every 14 days. Before, reimbursement prices were set every 6 months
Base: Working group	Sep. 15 2003 Jun. 07 2004	The Danish Ministry of Health announces to assemble a working group that is asked to submit proposals regarding reimbursement rules with the aim to increase competition. The Association of Danish Pharmacies launches the idea that reimbursements should be based on the cheapest domestic product within substitute groups. The idea earns widespread support among leading politicians
Announcement	Jun. 21 2004 Mar. 28 2005	The law regarding the new reimbursement system is passed by the Danish parliament
Treatment: Implementation	Apr. 01 2005 Sep. 25 2006	The new law is implemented
New LIF agreement	since Oct. 29 2006	The LIF and the government agree upon on a price ceiling corresponding to the price on 30 Aug. 2006

Appendix B: characterization of statins in terms of their ATC code

2-Level	3-Level	4-Level	5 - Level	
C10 Lipid Modifying Agents	C10A	C10AA HMG CoA reductase inhibitors (Statins)	C10AA01	**simvastatin**
			C10AA02	**lovastatin**
			C10AA03	**pravastatin**
			C10AA04	**fluvastatin**
			C10AA05	**atorvastatin**
			C10AA06	cerivastatin
			C10AA07	**rosuvastatin**
			C10AA08	pitavastatin
		C10AB Fibrates	C10AB01	clofibrate
			C10AB02	bezafibrate
			C10AB03	aluminium clofibrate
			C10AB04	**gemfibrozil**
			C10AB05	fenofibrate
			C10AB06	simfibrate
			C10AB07	ronifibrate
			C10AB08	ciprofibrate
			C10AB09	etofibrate
			C10AB10	clofibride
		C10AC Bile acid sequestrants	C10AC01	**colestyramine**
			C10AC02	**colestipol**
			C10AC03	colextran
			C10AC04	**colesevelam**
		C10AD Nicotinic acid and derivatives	C10AD01	niceritrol
			C10AD02	nicotinic acid
			C10AD03	nicofuranose
			C10AD04	aluminium nicotinate
			C10AD05	nicotinyl alcohol (pyridylcarbinol)
			C10AD06	**acipimox**
			C10AD52	nicotinic acid, combinations
		C10AX Other lipid modifying agents	C10AX01	dextrothyroxine
			C10AX02	probucol
			C10AX03	tiadenol
			C10AX05	meglutol
			C10AX06	omega-3-triglycerides incl. other esters and acids
			C10AX07	magnesium pyridoxal 5-phosphate glutamate
			C10AX08	policosanol
			C10AX09	**ezetimibe**
			C10AX10	alipogene tiparvovec
	C10B	C10BA combinations	C10BA01	lovastatin and nicotinic acid
			C10BA02	**simvastatin and ezetimibe**
		C10BX combinations	C10BX01	simvastatin and acetylsalicylic acid
			C10BX02	pravastatin and acetylsalicylic acid
			C10BX03	atorvastatin and amlodipine

Appendix B displays a detailed classification of lipid modifying agents with their respective ATC codes. Only boldfaced chemical substances are marketed in Denmark. Source: WHO Collaborating Centre for Drug Statistics Methodology.

Appendix C: Market and product characteristics

	All			Brand			Generics			Parallel Imports		
	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.	Mean	Median	Std. Dev.
All periods												
Number of products	122.27	127	14.16	34.42	35	3.79	60.81	60	11.73	32.08	29	9.41
Number of firms	20.00	19	3.12	4.92	5	0.27	9.95	10	1.71	5.64	5	1.75
Package size	67.75	98	35.79	67.84	98	34.77	67.61	98	37.16	67.93	84	34.17
Strength in mg.	28.62	20	18.56	33.92	20	23.67	27.77	20	15.99	24.09	20	14.60
Obs.	13'861			3'907			6'633			3'321		
Base period												
Number of products	126.39	127	4.85	37.23	38	0.97	54.59	55	2.24	35.28	35	5.24
Number of firms	19.01	19	0.44	5	5	0	10.01	10	0.44	5	5	0
Package size	65.97	98	34.96	64.62	98	35.06	66.24	98	35.04	66.98	84	34.72
Strength in mg.	26.70	20	16.55	34.35	20	22.85	23.85	20	12.20	22.93	20	10.70
Obs.	2'524			744			1'090			690		
Implementation period												
Number of products	126.03	132	15.07	33.70	35	2.57	70.34	72	7.24	23.00	26	6.01
Number of firms	22.17	23	2.30	5	5	0	10.83	11	1.22	7.07	8	1.48
Package size	67.49	98	34.69	69.87	98	34.47	66.75	98	35.13	66.17	60	33.43
Strength in mg.	30.70	20	19.97	33.41	20	24.19	29.79	20	17.68	29.38	20	19.31
Obs.	4'963			1'340			2'781			842		

Prices per Defined Daily Dose (DDD) in June 2005 DKK, where 1 DKK = 0.165 US dollars. Quantities measure as sold DDD.

Appendix D: Identification and estimation of the demand model

In the discussion of identification we closely follow recent propositions in Reynaert and Verboven (2012) about the benefits of using optimal instruments in random coefficient logit models. Subsequently, we sketch out our estimation procedure.

The unobserved characteristics of product j, ξ_{jt}, are known to both producers and patients, which implies that prices are endogenous in equilibrium and must be instrumented. Not instrumenting prices leads to downward biased estimates of the price coefficient α_i. We take two steps to remedy the problem of price endogeneity. First, we employ product name fixed effects to control for time-invariant quality levels. Second, a set of time-varying instruments accounts for variation around time-invariant means. Hence, identification relies on the conditional moment restrictions

$$E\left[\xi_{jt}|\boldsymbol{X_t}\right] = 0, \tag{8}$$

which is the mean independence of unobserved product quality ξ_{jt} of observed product characteristics $\boldsymbol{X}$.

These conditional moment restrictions can be transformed into unconditional moment restrictions

$$E\left[\xi_{jt}\boldsymbol{z_{jt}}\right] = 0, \tag{9}$$

where $\boldsymbol{z_{jt}}$ are the instruments. Reynaert and Verboven (2012) have shown that Chamberlain's (1987) optimal instruments work extremely well in random coefficient logit models, most importantly in identifying the nonlinear parameters. The set of optimal instruments is defined as the set of derivatives of the unobserved characteristic with respect the estimated parameters:

$$\boldsymbol{z_{jt}} = E\left[\left.\frac{\partial \xi_{jt}(\theta)}{\partial \theta'}\right| \boldsymbol{X_t}, \boldsymbol{w_{jt}}\right], \tag{10}$$

where we include an input price index as cost shifter w_{jt}. The intuition is equivalent to standard instruments with the difference that the derivatives make use of the functional forms assumed in the model whereas the standard instruments are simple linear projections. To see this, Reynaert and Verboven (2012) show that the set of derivatives with respect to the linear parameters β and α are simply the set of observed product characteristics and cost shifters. The derivative with respect to the nonlinear parameter σ is a nonlinear function of all competing products' characteristics. Hence, the biggest gain is achieved for the nonlinear parameter σ since the market share equation taking into account consumer heterogeneity can be exploited. Berry et al. (1999) and Goeree (2008) have previously approximated the expectation in equation (10) to construct optimal instruments. They evaluate the derivative at the mean of the disturbance vector (that is at $\xi_{jt} = 0$) while Reynaert and Verboven (2012) form the exact expectation by computing the mean of the derivative over $\hat{\xi}_{jt}$. The latter is the approach we follow.

Note that in order to compute $\boldsymbol{z_{jt}}$ in equation (10), we require initial estimates for $\boldsymbol{\theta}$ the very parameter vector we aim to ultimately estimate. One option would be to estimate the computationally expensive heterogenous logit model using standard instruments and using the results obtain therein as initial estimates for the optimal instruments. Reynaert and Verboven (2012) propose a simpler approach and show that it performs equally well as running the more general model twice. The idea is to estimate a homogenous IV logit model first. This is a linear IV regression and, hence, very fast. We choose three sets of standard instruments for this preliminary estimation. First, the sums of own other products' observed characteristics and sums of other firms' product characteristics which follows the arguments in Bresnahan (1987) and Berry et al. (1995) that the crowdedness in characteristics space should have an impact on equilibrium markups. Second, we include own product characteristics which are assumed to be exogenous. We follow Dubé et al. (2012) by also including squared and interaction terms of the product characteristics active ingredient

strength and package size. Third, we make use of a cost-side variable to account mainly for packaging costs. We interact an index for pulp and paper prices with product name fixed effects. This model does not obtain an estimate for σ so we must guess an initial value. We set this value equal to the absolute mean price coefficient $|\alpha|$. With these initial estimates at hand we can now compute the complete set of optimal instruments $\boldsymbol{z_{jt}}$ in equation (10).[9]

We estimate the random coefficient logit model using a sample that includes all products marketed between February 2003 and June 2007. In this sample, we observe 115 bi-weekly time-periods and approximately 100 products per period. Using our optimal instruments, we estimate the model by solving a mathematical program with equilibrium constraints (MPEC) as introduced by Su and Judd (2012) and Dubé et al. (2012):

$$\begin{aligned} \min_{\theta,\xi} \quad & \boldsymbol{g}(\boldsymbol{\xi})' \, \boldsymbol{W} \, \boldsymbol{g}(\boldsymbol{\xi}) \\ & \text{subject to} \quad \boldsymbol{s}(\boldsymbol{\xi};\boldsymbol{\theta}) = \boldsymbol{S}, \end{aligned}$$

where $\boldsymbol{g}(\boldsymbol{\xi})$ is the sample analogue to $E(\boldsymbol{z_{jt}}\boldsymbol{\xi})$. The main advantage of this approach as compared to the nested fixed point algorithm in Berry et al. (1995) is that the first and second derivatives of this problem are highly sparse in cases with many markets and not too many products. This can be exploited by numerical solvers and substantially increase computational speed. It also avoids numerical error propagation by circumventing the nesting of loops for optimization. We adapt and use Matlab code provided online by Dubé et al. (2012).

To obtain the constraints $\boldsymbol{s}(\boldsymbol{\xi};\boldsymbol{\theta}) = \boldsymbol{S}$ we solve the market share equation in (3) numerically. We assume ν to follow a standard normal distribution and draw 5000 modified latin hypercube sampling draws for estimation, as proposed in Hess et al. (2006), which have shown to be an improvement over frequently used Halton draws.[10] We further follow the proposition in Knittel and Metaxoglu (2012) to use 50 different starting values to increase confidence that the numerical solver stops at the true solution. The majority out of these 50 estimation runs converge, and those that do, converge to the same solution. The Knitro 8.0 solver's exit flag confirms convergence (as opposed to pre-mature stopping).

We compute changes in Marshallian consumer surplus. Our assumption of linear utility implies the absence of income effects so that consumer surplus and compensating variation coincide. The absence of income effects is a reasonable assumption if the change in consumer surplus is small relative to household income. This is the case for the Danish reform in the market for statins.

[9] See page 10 in Reynaert and Verboven (2012) for the exact algorithm we use.

[10] Consumer demographics such as the income distribution in Denmark are likely to explain some of this unobserved heterogeneity with respect to price sensitivities. Ideally, we would include the income distribution when estimating the distribution parameters for the price coefficient. However, given the shortness of the analyzed time period we do not observe much variation in the national income distribution in Denmark and, hence, including it will not lead to improved identification of the model. Furthermore, the fact that Denmark has a comparatively flat income distribution reduces the potential of including this observed consumer demographic.

Appendix E: First stage results for IV Logit specification

Strength of other firms' products	.0003***
	(.00007)
Strength of own products	-.0007***
	(.0002)
Strength	-.067***
	(.005)
Package size	-.017***
	(.002)
Strength^2	-.0001*
	(.00006)
Package size^2	.0001***
	(.00001)
Strength × package size	.0001***
	(.00004)

Dummy variables

Atorvastatin Ranbaxy	-99.24	Pravastatin Sandoz	-16.30***	Zarator	-13.60***
	(94.84)		(3.493)		(2.936)
Canef	-13.55***	Pravastatin Stada	-44.58***	Zocolip	-12.23***
	(3.264)		(16.06)		(3.370)
Crestor	-8.48***	Simvacop	-31.33***	January	.03
	(2.776)		(9.030)		(.111)
Lescol	-5.76**	Simvastatin 1A Farma	8.57**	February	.14
	(2.835)		(3.935)		(.122)
Lescol depot	-4.77*	Simvastatin Actavis	-6.09**	March	-.08
	(2.810)		(3.040)		(.120)
Lipitor	-8.91***	Simvastatin Alpharma	-2.03	April	-.06
	(2.942)		(2.978)		(.123)
Lovacodan	-4.21	Simvastatin Alternova	-4.78	May	-.28**
	(3.053)		(2.945)		(.124)
Lovastatin Actavis	-5.13*	Simvastatin Arrow	-.46	June	-.19
	(2.839)		(3.261)		(.124)
Lovastatin Alternova	-2.52	Simvastatin Genthon	-10.14	July	-.04
	(2.875)		(11.07)		(.134)
Lovastatin Universal Farma	-9.83***	Simvastatin Gevita	-5.59	August	-.03
	(2.849)		(3.936)		(.125)
Lovastatin ratiopharm	.82	Simvastatin Hexal	-9.87***	September	-.12
	(4.928)		(2.917)		(.128)
Mevacor	-15.58**	Simvastatin Merck NM	-19.15***	October	-.19
	(7.595)		(7.145)		(.121)
Perichol	-5.00	Simvastatin Orifarm	-24.96**	November	-.16
	(3.422)		(10.62)		(.117)
Pravachol	-38.73***	Simvastatin Paranova	-78.24***	LIF Agreement	-2.23***
	(3.583)		(14.20)		(.241)
Pravastatin 1A Farma	-14.57***	Simvastatin Ratiopharm	-25.34***	Adjustment	-.78***
	(3.533)		(4.482)		(.211)
Pravastatin Alternova	-4.36	Simvastatin Sandoz	-12.50***	Working group	-.48***
	(3.260)		(3.025)		(.165)
Pravastatin HEXAL	-19.09***	Sortis	-3.10	Announcement	-.76***
	(3.529)		(3.884)		(.152)
Pravastatin Nycomed	-9.62***	Statinacop	-2.98	Implementation	-.13
	(2.858)		(4.002)		(.133)
Pravastatin Ranbaxy	-15.39***	Tahor	-7.00**	P&P × Name	Yes
	(3.875)		(3.384)		
Pravastatin Recept	-9.35**	Torvast	-12.49***	Constant	12.44***
	(4.083)		(3.822)		(3.156)

F-test results	
All instruments	136.16
BLP instruments	30.91
Pulp & Paper (P&P) instruments	16.35
Squares and interactions instruments	13.70
R^2	.59

Notes: First stage regression coefficients of IV Logit. Robust standard errors in parentheses. The reference categories for dummy variables are: Product name Zocor, Month December, and Period 'New LIF agreement'.
*** p< 0.01, ** p< 0.05, * p < 0.1

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