

Consumer Information and Pharmaceutical Prices: Theory and Evidence*

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Abstract

The impact of a reform that increased consumer information on brand name and generic pharmaceutical prices is analyzed both theoretically and empirically. The theoretical results show that an increase in information likely reduces the price of brand name pharmaceuticals, while the results regarding generics are less clear. In the empirical part of the paper, the introduction of the substitution reform in the Swedish pharmaceuticals market in October 2002 is used as a natural experiment regarding the effects of increased consumer information on pharmaceutical prices. The results clearly show that the reform has lowered the price of both brand name and generic pharmaceuticals.

Key Words: pharmaceutical industry; generic competition; generic drugs; brand name drugs

JEL classification: D80; D83; L65; I11

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1 Introduction

In this paper, a reform of the Swedish pharmaceutical market will be used to study the effects of increased consumer information about price differences in medically equivalent treatments (brand name and generic drugs) on the pricing policy of pharmaceutical firms.

There is a vast theoretical literature concerning the effects of imperfect consumer information on pricing and market structure. Stigler (1961) showed that imperfect information creates market power and prices above competitive levels; while Diamond (1971) showed that if information is costly, this could lead to an equilibrium where firms charged the monopoly price rather than the competitive price. In addition, Salop and Stiglitz (1977) presented a model where low-cost stores had higher sales because low search cost individuals actively seek them out, while only high search cost individuals patronized the high cost stores. They also established that costly information could lead to market equilibria where the law of one price does not hold. In a related study, Stiglitz (1979) showed that under imperfect information, increasing the number of firms could actually reduce competition and increase prices.

Turning to theoretical models of pharmaceutical markets, Frank and Salkever (1992) have shown that an increase in the number of informed consumers will lead to lower brand name pharmaceutical prices. Also, based on a conjecture by Grabowski and Vernon (1992) that an increase in the number of informed customers (i.e. cross-price sensitive consumers) would increase the negative impact of entry on brand name pharmaceutical prices, Frank and Salkever showed that, theoretically, this is not necessarily the case. Our theoretical model, presented in section 3 below, will be a modified version of the Frank and Salkever model.

Only a few empirical tests of how consumer information affects prices in health care markets exist. Pauly and Satterthwaite (1981) studied how the market for a reputation good (i.e. a good that is marketed mainly through recommendation from friends, relatives or colleagues) such as physician services was affected by additional primary care physicians entering the market. Since this increases consumer search costs in their model, this will also make each primary care physicians demand curve become less elastic, leading to higher prices. In a more recent paper, Sorensen (2000) studied the relationship between imperfect consumer information and prices among prescription pharmaceuticals. The

data were collected from pharmacies in upstate New York, and the price differences among medically equivalent prescriptions were found to be large. In fact, on average the highest listed price exceeded the cheapest available alternative by as much as 50 percent. In addition, the results give support to a consumer search cost model, since frequently purchased pharmaceuticals had both lower markups and lower price differences when compared to one-time prescription pharmaceuticals.

In October 2002, the Swedish pharmaceuticals market was reformed. The reform (Lag 2002:160) required that pharmacists inform the consumers if there are substitute products available, as well as that the cheapest available substitute product would be provided within the Swedish pharmaceuticals insurance system. The reform also required that the consumers are given the opportunity to buy the prescribed pharmaceutical product instead of the cheapest available product, paying the difference in price between the products themselves. This means that under the new regulations, consumers have more information about the price difference between the prescribed (often brand name) product and the cheapest available (generic) alternative than they had prior to the reform.

There were some other changes due to the reform that could also have affected the pricing decision of pharmaceutical firms. First, the out-of-pocket cost for patients changed somewhat. Second, the transaction cost of generic substitution was lowered when the reform was introduced. In this paper we will argue that although all three changes could have affected the market, the most important change was the increase in consumer information. Thus, the introduction of the substitution reform will be used to empirically measure the effects of increased consumer information on brand name and generic pharmaceutical prices.

Our theoretical model shows that an increase in information is likely to reduce the price of brand name pharmaceuticals, but the results regarding generics are less clear. Hence, the main hypothesis to be tested is if the substitution reform, by increasing consumer information about pharmaceutical prices, has decreased the price of brand name and/or generic pharmaceuticals. In addition, we will test whether the possible price response differs between brand name and generic drugs and also study additional heterogeneity in the reform-effect, suggested by the theoretical model. The empirical analysis is performed using monthly data on pharmaceuticals sold January 2001 to October 2006.

The paper contributes to the existing literature in the following ways; first, contrary to previous theoretical studies, our theoretical model analyses how both brand name and generic pharmaceuticals are affected by increased consumer information. Second, previous studies of the introduction of substitution reforms in European pharmaceuticals markets (e.g. Buzzelli et al., 2006) use pharmaceutical price indexes as their dependent variable. As such, they cannot study heterogeneity in the reform-effect among brand name and generic drugs. The introduction of the substitution reform and the use of individual pharmaceutical price data in our paper make it possible to study both how pharmaceutical prices were affected by the reform, as well as possible heterogeneity in the reform-effect among brand name pharmaceuticals and generics. The main finding from this paper is that the reform has lowered the price in both pharmaceutical groups.

The next section describes the substitution reform, while section 3 presents the theoretical model, based on Frank and Salkever (1992). Section 4 presents the data and the empirical model to be used in this study, as well as the results from the estimations. Finally, in section 5 the paper's conclusions are presented.

2 The substitution reform

The substitution reform came into effect on October 1, 2002. As mentioned above, the reform required that pharmacists inform the consumers if there are substitute products available, as well as that the cheapest available generic substitute product (which is considered to be a perfect substitute for the brand name drug by the Swedish Medical Products Agency) would be provided within the Swedish pharmaceuticals insurance system. The pharmacist must also inform the consumers that they can buy the prescribed pharmaceutical product instead of the generic if they pay the difference in price between the products themselves. As such, the new regulations provide consumers with more information about the price difference between the prescribed product and the cheapest available generic alternative than before. Finally, the reform requires that pharmacists substitute the prescribed pharmaceutical product to the cheapest available generic (or parallel imported product) in cases when neither the prescribing physician prohibits the switch for medical reasons, nor the consumer chooses to pay the price difference between the prescribed and the generic alter-

native. In cases where the physician prohibits the switch due to medical reasons the consumer is still reimbursed.

The reform was supposed to lower pharmaceutical costs in two different ways; directly, as more expensive pharmaceuticals were exchanged for cheaper generic copies, and indirectly through increased price competition. The latter effect will be studied in this paper.

Pharmaceutical firms decide which prices they charge for pharmaceuticals in Sweden, but for products to be included in the Swedish pharmaceuticals insurance system the price charged by the pharmaceutical firms has to be authorized by the Pharmaceutical Benefits Board. Pharmaceuticals are sold through a nation wide government owned monopoly, the National Corporation of Swedish Pharmacies (NCSP), which has a margin on the pharmaceutical products that is determined by the Pharmaceutical Benefits Board. The regulations also imply that the NCSP is required to charge a nation wide uniform price for each pharmaceutical product in Sweden.

Before the substitution reform, a reference price system introduced in January 1993 was in effect.¹ Under that system, the Swedish National Social Insurance Board set a reference price equal to 110 percent of the price of the cheapest available generic product, and all costs exceeding this reference price were to be borne by the consumer (RFFS 1992:20, 1996:31).

Some other changes due to the reform might also have affected the pricing policies of pharmaceutical firms. First, the out-of-pocket cost for patients changed when the reference price was abolished. Under the substitution reform costs up to 100 percent of the cheapest generic alternative is included in the pharmaceutical insurance system, compared to 110 percent during the reference price system. This increased the patients out-of-pocket costs for choosing to buy the prescribed pharmaceutical with 0-10 percent of the price of the cheapest generic version, depending on the patient's co-payment rate in the insurance system. On average this means an extra out-of-pocket cost of approximately 19 SEK (\approx 2 EURO).² As such, the change in consumer out-of-pocket payments

¹The effects of the reference price system on pharmaceutical prices have been analyzed previously, see e.g. Aronsson et al. (2001) and Bergman and Rudholm (2003).

²The calculation is based on the fact that the average price of the prescribed products and the available substitute products in the substitution system was approximately 300 SEK and 250 SEK, respectively, and a patient co-payment rate of 25 percent. $9.51 \text{ SEK} = 1 \text{ EURO}$, exchange rate 2008-09-12.

due to the reform was quite small. Second, even before the reform a prescribed pharmaceutical product could be substituted for a cheaper generic version if the prescribing physician had given his/her consent to this on the prescription, or if the patient requested substitution. In the latter case there was, however, a recommendation that the physician be contacted before substituting products if possible, and a requirement that the prescribing physician be informed about the substitution after. This means that the transaction cost of generic substitution was lowered when the reform was introduced in 2002, which could also affect the pricing behavior of pharmaceutical firms. However, physicians that did not want to be disturbed by pharmacy personnel wanting to make generic substitutions should have consented to generic substitution even before the reform.³ Thus, we argue that the main reform effect should be due to the increase in consumer information.

3 Theoretical model

To study the effects of increased information on brand name and generic drugs, and to provide an theoretic underpinning for studying heterogeneity in these effects, we turn to the model developed by Frank and Salkever (1992). We modify their model to be able to analyze the price of generics as well. In a specific market there is one brand name firm, b , and $n - 1$ generic firms, that all produce one pharmaceutical product each and have identical cost functions. Since there are consumers who are willing to pay the difference in price between the brand name and generic pharmaceutical in order to get the brand name pharmaceutical, it will be assumed that consumers actually view the brand name and the different generic drugs as close substitutes, instead of perfect ones.⁴ Therefore, a firm cannot get all informed costumers by setting their price slightly below the others. The consumers only differ in their preferences for the drugs and the information they have about prices.

³We have not been able to find any studies of how common it was for prescribing physicians to allow generic substitution before the reform. However, during the first 15 months after the substitution reform, physicians chose to allow the exchange in 97 percent of the cases (National Corporation of Swedish Pharmacies et al., 2004).

⁴During the first 15 months after the reform, consumers chose to deny the exchange in approximately 5 percent of the cases (National Corporation of Swedish Pharmacies et al., 2004).

Several studies have reported that brand name products, in spite of considerably higher prices than the available generic substitutes, are able to maintain dominant positions in terms of market shares (e.g. Bond and Lean, 1977; Statman, 1981; Hurwitz and Caves, 1988). For the Swedish pharmaceuticals market, Aronsson et al. (2001) reported that some of the products in their study maintained market shares in excess of 80 percent even though they had a more than 50 percent markup over the price of the available generic substitutes. As such, the principal difference between a generic firm and the brand name firm in our theoretical model is that the brand name firm will be assumed to have a higher market share compared to the generic firm at equal prices, due to consumer preferences for the brand name product. Due to its high market share, we assume that a change in the brand name price will have a significant impact on the generic prices and treat the brand name firm as a Stackelberg leader which incorporates the generic firms' price-responses in its pricing decision.⁵ It will also be assumed that the generic firms take all other prices as given in the sense that they do not take into account how their pricing decisions will affect that of other pharmaceutical firms. As such, the analysis will start by analyzing the pricing decision made by the generic firms.

The demand function for generic product i is written

$$Q_i = \alpha D_i^I(P_i, P_j, P_b, n) + (1 - \alpha) D_i^{UI}(P_i, n). \quad (1)$$

α is the share of informed consumers in the market, who know the prices of all available pharmaceutical products, and $(1 - \alpha)$ is the share of uninformed consumers in the market, who only know the price of the drug they are buying. αD_i^I and $(1 - \alpha) D_i^{UI}$ represent the demand facing the generic firm i from informed and uninformed consumers, respectively. P_i, P_j and P_b are the prices of the generic drug i , all other generic drugs and the brand name drug, respectively, which together with the number of pharmaceutical products, n , affect the demand for the generic drug i from the informed consumers. The demand from the uninformed consumers is not affected by other prices, but by the number of pharmaceutical products.

⁵The qualitative results from the theoretical model hold even if the brand name firm is treated as a Nash player towards the generic ones.

The generic firm i 's profit function is written

$$\begin{aligned} \pi_i = & P_i * [\alpha D_i^I(P_i, P_j, P_b, n) + (1 - \alpha) D_i^{UI}(P_i, n)] \\ & - C_i(\alpha D_i^I(P_i, P_j, P_b, n) + (1 - \alpha) D_i^{UI}(P_i, n)), \end{aligned} \quad (2)$$

where $C_i(Q_i)$ is the firm's cost function. The marginal cost is assumed to be positive and constant.⁶ The firm chooses P_i to maximize the profit, which gives the following first order condition

$$\begin{aligned} d\pi_i/dP_i = & \left(P_i - \frac{dC_i}{dQ_i} \right) * \left[\alpha \frac{dD_i^I}{dP_i} + (1 - \alpha) \frac{dD_i^{UI}}{dP_i} \right] \\ & + \alpha D_i^I(P_i, P_j, P_b, n) + (1 - \alpha) D_i^{UI}(P_i, n) = 0. \end{aligned} \quad (3)$$

The price function of generic product i can thus be written $P_i(\alpha, n, P_j, P_b)$. All generic firms are assumed to face identical demand and cost functions and will therefore set the same price, which we denote by P_g . The best response function of generics is written $P_g(\alpha, n, P_b)$.

Accordingly, the brand name producer's demand function is written

$$Q_b = \alpha D_b^I(P_b, P_g(\alpha, n, P_b)) + (1 - \alpha) D_b^{UI}(P_b, n), \quad (4)$$

which gives the following first order condition

$$\begin{aligned} d\pi_b/dP_b = & \left(P_b - \frac{dC_b}{dQ_b} \right) * \left[\alpha \left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g} \frac{\partial P_g}{\partial P_b} \right) + (1 - \alpha) \frac{dD_b^{UI}}{dP_b} \right] \\ & + \alpha D_b^I(P_b, P_g(\alpha, n, P_b)) + (1 - \alpha) D_b^{UI}(P_b, n) = 0. \end{aligned} \quad (5)$$

Equations (3) and (5) differ from each other since the brand name firm get higher sales compared to a generic firm with the same price and because the brand name firm act as a Stackelberg leader. If $\frac{\partial P_g}{\partial P_b} > 0$,⁷ which we assume, either one of these differences guarantee that the brand name firm will set a higher price than the generic ones.

⁶This assumption is not crucial for the qualitative results from the model, but it simplifies expressions and makes them easier to interpret.

⁷See Appendix 1 for a discussion of these assumptions.

3.1 The effect of an increase in information

The effect of an increase in the share of informed consumers (i.e. an increase in α) on the price of generic drug i is written

$$dP_i/d\alpha = \left\{ \left(P_i - \frac{dC_i}{dQ_i} \right) * \left(\frac{dD_i^I}{dP_i} - \frac{dD_i^{UI}}{dP_i} \right) + (D_i^I - D_i^{UI}) + \frac{dP_i}{dP_b} \frac{dP_b}{d\alpha} + \frac{dP_i}{dP_j} \frac{dP_j}{d\alpha} \right\} / (-\delta_i) \quad (6)$$

We assume that the second order sufficient condition for a maximum, $\delta_i < 0$, is fulfilled (Appendix 2). The mark-up, $P_i - \frac{dC_i}{dQ_i}$, is positive. The second term in the numerator, and therefore the first product, will be negative if the own price response of informed consumers is greater than the own price response of uninformed consumers, which is a reasonable assumption. The next two terms show how the total demand for generic i is affected by a change in the share of informed consumers. These terms are jointly positive since informed consumers are more likely to patronize the generic product i due to price information, while consumers in the uninformed group choose the pharmaceutical product closest to their personal preferences, brand name or generic.

The final terms show the indirect effect on P_i of an increase in α , working through the effect of α on other prices. The terms $\frac{dP_i}{dP_j}$ and $\frac{dP_i}{dP_b}$, which are discussed in more detail in Appendix 1, are assumed to be positive. The last two terms therefore imply that a generic firm is more likely to reduce its price if the brand name price is reduced and that any price change of a generic firm, due to symmetry, is enhanced by the price changes of the other generic firms.

To sum up, the direct effect of an increase in the share of informed consumers on the generic prices is not clear cut. The first product works in the direction of lower prices as long as the consumer becomes more price sensitive when information about prices increases. On the other hand, the next two terms in the numerator, D_i^I and D_i^{UI} , together have a positive effect on the price change for generic product i . Finally, the sign of the indirect effect depends on the effect of α on the brand name price, to which we now will turn.

The effect of an increase in α on the brand name price is written

$$dP_b/d\alpha = \left\{ \left(P_b - \frac{dC_b}{dQ_b} \right) * \left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g} \frac{\partial P_g}{\partial P_b} - \frac{dD_b^{UI}}{dP_b} \right) + (D_b^I - D_b^{UI}) + \frac{dP_b}{dP_g} \frac{dP_g}{d\alpha} \right\} / (-\delta_b). \quad (7)$$

$P_b - \frac{dC_b}{dQ_b}$ is positive and larger than the mark-up for a generic firm. The second term will be negative if the own price response of informed consumers is greater than the own price response of uninformed consumers, which seems reasonable to assume. The brand name firm's Stackelberg role, however, works for that this term will be less negative than the corresponding one for the generic firms. The next two terms show how the total demand for the brand name product is affected by a change in the share of informed consumers. Since informed consumers are aware of the cheaper generic substitutes, they buy less from the brand name firm than uninformed consumers and the sum of these terms will be negative. That is, the brand name firm, unlike the generic companies, loses sales if there is an increase in the share of informed consumers, which reduces the revenue loss from price reductions. Therefore, this works for a higher price reduction for brand name drugs compared to the generic drugs. The sign of the final term depends on the price change of the generic firms.

Given our assumptions, the theoretical model shows that it is likely that the substitution reform, which is our exogenous measure of increased consumer information about pharmaceutical prices, will lead to price reductions for brand name drugs while the predictions are less clear for the generic products. We now turn to the task of empirically testing the predictions of our model, i.e. that the prices of brand name (and perhaps also generic) pharmaceuticals will be reduced by the substitution reform.

4 The empirical analysis

4.1 Data and empirical specification

The pharmaceutical prices in Sweden during the period January 2001 to October 2006 are extracted from a dataset provided by the County Council of Västerbotten, Sweden. The data cover all prescription pharmaceuticals sold in the county during the period and sums up to a total number of 15 million

prescriptions.

Only observations referring to pharmaceuticals considered to be exchangeable within the substitution reform are relevant for this study. We identified these observations by a two step procedure. First, we used a wide definition of which pharmaceuticals that should be considered to be exchangeable within the substitution reform. This wide definition is based on the criteria specified by the Medical Product Agency, namely that the different pharmaceutical products have the same active substance, strength, and form (pills, oral fluids, etc.) and that the packages are of similar size. In a second step, these commodities were more carefully examined by manually comparing them to the Medical Product Agency list over substitutable products published at the time of the introduction of the reform. After this examination 4082 commodities, in 856 different exchange groups, are defined to be substitutes according to the regulations set out in the exchange reform. These commodities account for approximately 7 million prescriptions in our original sample. We can identify in which month each prescription is sold, and there are thus a maximum of 70 observations for each commodity, 21 before the reform and 49 after, giving a maximum number of observations of 285 740 ($=4082*70$). However, not all of these pharmaceutical products are sold each month, reducing the actual number of observations to 126 904. Due to missing data for one or more of our key variables, the final sample consists of 105 587 observations.

Our dataset does not include information about which product that is the brand name product.⁸ From the theoretical model we know that brand name firms will distinguish themselves by setting higher prices than their competitors. This is also supported by previous empirical studies (e.g. Aronsson et al., 2001; Bond and Lean, 1977; Statman, 1981; Hurwitz and Caves, 1988). Thus, *Brand* is an indicator variable that takes the value one if the pharmaceutical has had the highest price in the exchange group both in the first and third six months periods prior to the reform. The six month period is used to reduce the problem of not all commodities being sold each month, and two examination periods are used to reduce the risk of our labeling being affected by temporary price changes. We choose examination periods prior to the reform, since the

⁸Note that in our dataset, so called "me-to-pharmaceuticals" fall in their own exchange groups. This coincides with the regulations set out in the substitution reform, where substitution to "me-to-pharmaceuticals" is not allowed.

reform may reduce price differences between brand name and generic drugs, making labeling based on periods after the reform less robust. In the 9 percent of the cases where the pharmaceutical had the highest price in only one of these periods, a closer examination was performed where a pharmaceutical was classified as a brand name drug if there were only one producer not belonging to the group of the most common generic producers in the Swedish market.⁹ The remaining pharmaceuticals, where no apparent brand name producer could be identified were all treated as generics.^{10,11}

The main purpose of the empirical part of this paper is to study the effect of the introduction of the substitution reform on the prices of brand name and generic pharmaceuticals. Therefore, our focus will be to achieve unbiased estimates of the reform-effect. As such, we use product-specific fixed effects to capture several of the effects that according to equations (3) and (5) can explain the difference in price levels between pharmaceutical products, and focus our attention on changes in the prices caused by changes in the shares of informed consumers, as described in equations (6) and (7). The latter two equations reveal several variables which will influence the reform-effect. We will therefore take them as starting points to study heterogeneity in the reform-effect as well. Since we are not able to model all differences between the two equations, we estimate the models separately for brand name products and generics, allowing the parameter estimates to differ. However, the specification for the two groups will include the same variables. Below, we therefore only discuss the

⁹In Sweden, there are some producers that specialize in producing and selling generic products in several of the substitution groups covered in this study. These are easily identified in our dataset.

¹⁰All regressions have also been performed excluding these 9 percent of the cases. All qualitative results from the estimations presented below are unchanged, and these results have thus been excluded in order to save space.

¹¹To test whether this method can identify the correct brand name pharmaceuticals a dataset from Rudholm (2001) is used. In that dataset, the brand name pharmaceuticals were identified for 22 substances. Comparing the brand name pharmaceuticals in that dataset to ours, 5 of the substances have been removed from the market. Of the 17 remaining substances (corresponding to 12 percent of the prescriptions in our sample), the brand name product coincide for all product types and package sizes in 13 cases. In the remaining 4 cases, the brand name products coincide for tablets and capsules, but not for other types of products (e.g. oral fluids, intravenous fluids, etc). Such products correspond to approximately 14 percent of the observations in our sample. As such, we have re-estimated the empirical models using a dataset including only tablets and capsules. All qualitative results from these estimations coincide with the ones presented below. The results are available from the authors on request.

specifications for one of the two groups, brand name products.

From equation (7) we know that the effect of an increase in α on the brand name price can be written

$$\begin{aligned} dP_b = & \left(P_b - \frac{dC_b}{dQ_b} \right) * \left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g} \frac{\partial P_g}{\partial P_b} - \frac{dD_b^{UI}}{dP_b} \right) / (-\delta_b) d\alpha \\ & + (D_b^I - D_b^{UI}) / (-\delta_b) d\alpha + \frac{dP_b}{dP_g} \frac{dP_g}{d\alpha} / (-\delta_b) d\alpha. \end{aligned} \quad (8)$$

The first product on the right hand side includes the following four terms: the markup over marginal cost, the difference in own price response of informed and uninformed consumers, the second derivative of the brand name firm's profit maximization problem, and the change in information due to the introduction of the substitution reform. Our data allow us to study heterogeneity in the first and fourth terms.¹² First, the markup is assumed to be a linear function of our proxy variable $\ln markup_i$, that is

$$\left(P_b - \frac{dC_b}{dQ_b} \right) = \gamma_0 + \gamma_1 \ln markup_i. \quad (9)$$

$\ln markup_i$ is the logarithm of the difference between the average price for pharmaceutical product i in the pre-reform period and the minimum price in the group of exchangeable products during that period.¹³ As such, this variable is defined for pharmaceutical i . Mathematical formulas used to calculate this and other proxy variables are presented in Appendix 3. Second, the change in information is written

$$d\alpha = \gamma_3 dD_t + \gamma_4 d(\text{info}_g * D_t), \quad (10)$$

where D is an indicator variable taking the value one after the introduction of the substitution reform, and info_g is our proxy for the level of information available for consumers in exchange group g before the introduction of the reform.

¹²We do not model any heterogeneity in the second and third term. This reduces the number of interaction terms in our full model, but, similarly to the omission of interaction terms later discussed in Appendix 5, this does not cause any bias in the estimator of the reform effect.

¹³Using the minimum price in the group of exchangeable products as a proxy for marginal cost is similar to the approaches adopted by Grabowski and Vernon (1992) and Rudholm (2001). The logarithm is used since the dependent variable will also be in logarithmic form. γ_0 is included to capture the average of the difference between the minimum price and the marginal cost.

This is defined as the extra expenditure on pharmaceuticals in exchange group g prior to the reform compared to the expenditure if the given quantity would have been bought at the lowest price in the exchange group each month, measured in percent. Following Sorensen (2000), this measure is based on the notion that if consumers systematically and over time pay more than necessary for a given pharmaceutical within an exchange group, this reflects that consumers in that group are less informed about prices of available generic substitutes than in other groups, *cet. par.* The idea is that the change in the share of informed consumers will be larger in markets with low initial information. Multiplication gives the following expression for the first product in equation (8)

$$\begin{aligned} \left(P_b - \frac{dC_b}{dQ_b}\right) * d\alpha &= \gamma_0\gamma_3 dD_t + \gamma_0\gamma_4 d(\text{info}_g * D_t) \\ &+ \gamma_1\gamma_3 (\ln \text{markup}_i * dD_t) \\ &+ \gamma_1\gamma_4 [\ln \text{markup}_i * d(\text{info}_g * D_t)]. \end{aligned} \quad (11)$$

In the second product in equation (8), our division of the sample in brand name and generic pharmaceuticals in part controls for differences in the term $D^I - D^{UI}$ between pharmaceutical firms, and we do not model any remaining heterogeneity in this term or in the second-order derivative, δ_b . What remains in the second product is $d\alpha$, which is measured as described above.

The third product in equation (8) shows that the change in the price of the brand name producer can also be affected by price changes of generic products, caused by a change in the proportion of informed consumers. The effect on the generic prices in an exchange group mostly depends on the same variables as the effect on the brand name prices. In addition it depends on the markup of the generic firms, which does not directly affect the price change of the brand name drug. To study whether this indirect effect has any significant effect on the brand name price, we therefore identify the term $\frac{dP_g}{d\alpha}$ by $\text{diff} \ln \text{markup}_i$, which is the difference between the average markup in the exchange group g and the markup for the brand name drug i .¹⁴ $d\alpha$ is measured as presented above and we do not model any heterogeneity in the term $\frac{dP_b}{dP_g}$. Therefore, the third product in equation (8) leaves us with the two additional interaction terms $\text{diff} \ln \text{markup} * D$ and $\text{diff} \ln \text{markup} * \text{info} * D$.

¹⁴This specification for the logarithmic markup of the other drugs in the exchange group is chosen in order to reduce the multicollinearity problem which otherwise could be a problem.

So far, only the variables of main interest for this study as presented in equation (8) have been discussed, i.e. the variables that are supposed to capture the effects of increased consumer information due to the substitution reform. However, as can be seen in equation (5), the price will also depend on n , which in our empirical model will be measured by the number of products sold in a specific substitution group each month. The interaction terms $n * D$ and $n * info * D$ are included since the studies of Stiglitz (1979) and Frank and Salkever (1992) indicated that the effect on prices of the number of products can be affected by the share of informed consumers. A trend variable, $Trend$, is included in order to account for possible common price trends. Finally, product-specific fixed effects, θ_i , have been included. The empirical specification of the most general price equation for brand name and generic pharmaceuticals can thus be written

$$\begin{aligned}
\ln Price_{it} = & \beta_1 D_t + \beta_2 (info_g * D_t) + \beta_3 (\ln markup_i * D_t) \\
& + \beta_4 (\ln markup_i * info_g * D_t) + \beta_5 (diff \ln markup_i * D_t) \\
& + \beta_6 (diff \ln markup_i * info_g * D_t) + \beta_7 n_{it} + \beta_8 (n_{it} * D_t) \\
& + \beta_9 (n_{it} * info_g * D_t) + \beta_{10} Trend_t + \theta_i + \varepsilon_{it}. \tag{12}
\end{aligned}$$

Descriptive statistics for all variables included in the empirical estimations are presented in Table 1. $\ln Price$ is the logarithm of the price of the pharmaceutical per one hundred defined daily doses expressed in Swedish crowns in fixed October 2006 prices.¹⁵ Comparing the descriptive statistics for brand name and generic products, prices are roughly the same in the two groups. However, as can be seen in Table 1, the markup over marginal cost is larger for the sample containing brand name pharmaceuticals. Also, the descriptive statistics for $diff \ln markup$ show that the competing products in an exchange group on average have a lower price than the brand name product, while the opposite is true for generics. As such, the similar prices for brands and generics in Table 1 are due to a selection effect, where generic firms to a large extent have chosen to enter markets where the price is high. In addition, Table 1 shows that the number of products, n , is larger in the subsample for generics, as is $info$.

¹⁵ Defined daily doses (DDDs) is a World Health Organization measure of drug quantity.

Table 1 about here.

In equation (12), the following variables might be endogenous in the sense that these variables correlate with the error term; *info*, $\ln markup$, $diff \ln markup$ and n . The available dataset does not contain any reasonable variables to be used as instruments, hence endogeneity has to be addressed in some other way. In order to be able to later conclude which result that might be driven by endogeneity, we therefore discuss in Appendix 4 how the correlation between the error term and these variables are expected to affect the estimates of the different parameters. One should however keep in mind that these variables are proxy-variables and that the estimators can therefore also be biased due to measurement errors. Since bias in the estimators related to the variables mentioned above might affect the estimated reform-effect, we show in Appendix 5 how we can create a simple model, without the interaction variables, in order to estimate the reform-effect. In short, we show that the estimator for the reform parameter, β_1 , will be an unbiased estimator of the average reform-effect if the interaction variables are left out from the specification. The explanation is that this estimator will capture the linear relationship between this variable and the omitted interaction variables (e.g. Greene, 2003, Chapter 8).

Due to low variation in the number of products during the period of study, the remaining possibly endogenous variable, n , most likely only has a very limited influence on the estimated reform-effect. During the study-period, it is likely that the bulk of the correlation between n and the reform is due to the effect the reform has on the number of products, and that only a smaller fraction of the correlation is due to exogenous variation in n . Therefore, not controlling for n in the regression probably gives a better estimate of the total reform-effect.¹⁶ As such, the specification which will be used to estimate the average effect of the reform for both brand name and generic pharmaceutical subsamples is written

$$\ln Price_{it} = \beta_1 D_t + \beta_{10} Trend_t + \theta_i + \varepsilon_{it}, \quad (13)$$

where all included variables are exogenous. In all estimations, we use a Prais-Winsten estimator which corrects for first order serial correlation in the error

¹⁶We will also estimate the basic model including the number of firms, n , in order to study how this affects the parameter estimates for the reform indicator variable.

terms. In addition the error terms are allowed to be heteroskedastic and correlated within ATC-groups.¹⁷

4.2 The adjustment process

Two circumstances give firms an incentive to gradually adjust their price after the reform.¹⁸ First, there is incomplete information for firms about the reactions of consumers and other firms to the reform. Second, price increases are rarely allowed by the Pharmaceutical Benefits Board. To capture this, we introduce the variable $D/(t - R)$, where R is the time for the reform. We let the denominator be raised to the power μ , where μ is a parameter that will measure the curvature of the adjustment process. As such, our basic empirical specification including the adjustment process is written

$$\ln Price_{it} = \beta_1 D_t + \beta_{10} Trend_t + \gamma [D_t / (t - R)^\mu] + \theta_i + \varepsilon_{it}. \quad (14)$$

The models with adjustment are non-linear in the adjustment variable $D/(t - R)$. Since the models are nonlinear only due to one parameter, μ , it is convenient to estimate the models using a grid-search estimation strategy. We have employed this method for each model setting μ to values ranging from 0 to 4 and then estimating the remaining parameters using a Prais-Winsten estimator. Finally, likelihood values were used to discriminate between the different parameter values. The likelihood values were as well used to calculate 95 percent confidence intervals for the adjustment parameter, μ . As can be seen in the notes to Tables 2 and 3, the confidence intervals are not symmetric around

¹⁷In the World Health Organization's Anatomical Therapeutic Chemical (ATC) classification system, the drugs are divided into different groups according to the organ or system on which they act, and their chemical, pharmacological and therapeutic properties. In the ATC-groups used here, drugs which share the same chemical substances are grouped together.

¹⁸Since pharmaceutical firms knew about the reform before it came into effect in October, 2002, it is also possible that they started to adjust to the reform before this date. If this was the case, we would expect to obtain larger estimates of the reform effect if the estimations were performed excluding observations from the months directly before the reform. To study the importance of this possibility, we estimated all models excluding observations originating from April 2002, when the law regarding the reform was passed by parliament, until October 2002, as well as from January 2002, when the bill was presented to parliament, until October 2002, respectively. Although some of the point estimates in these regressions were somewhat larger than the ones presented in this paper, these differences were not statistically significant at conventional levels. Thus, these results indicate that the potential adjustments before the reform were too small to have any important impact on our estimates.

the point estimates. This is expected since a value of μ equaling zero lead to an empirical model where the adjustment variable becomes equal to the reform indicator variable.

4.3 Results

The results from the estimation for brand name products are presented in Table 2. In order to study heterogeneity in the reform-effect, we first discuss the results of the full model. For all variables that interact with other variables, we only discuss the differential and the derivatives, respectively, which are evaluated at the mean value for each variable. The estimate of the differential $\Delta \ln Price / \Delta D$ from the full model shows that the average reduction in the price of brand name pharmaceuticals due to the reform is approximately 2 percent,¹⁹ and there is a significant negative trend showing an average decrease in pharmaceutical prices of 0.2 percent per month.²⁰ From the theoretical model it is expected that if consumers had a low level of information prior to the reform (i.e. they systematically spent more money than necessary on the pharmaceuticals they bought), the reduction in brand name price after the introduction would be large. As can be seen in the lower half of Table 2, this is also the case. However, this result must be interpreted with caution since the estimators of the parameters for *info* may be negatively biased, as discussed in Appendix 4.

Table 2 about here.

¹⁹Since the dependent variable is in logarithmic form, the exact change in price (in percent) should for dummy variables be calculated using the formula $100 * [\exp(\beta) - 1]$. However, since the parameter estimates are small, using the exact formula gives the same results as using the parameter estimates directly after rounding.

²⁰Negative trend estimates are expected for two reasons. First, the price index for pharmaceuticals in Sweden has declined during the period of study. Second, since the price index also is affected by entry of new, relative expensive pharmaceuticals, the price of existing pharmaceuticals have on average a more negative development than the price index. Still, one might suspect that part of the reform effect is captured by the time trend. We therefore estimate the time trend for different time periods prior to the reform. Both for brands and generics, these estimates are slightly more negative compared to the estimated time trend for the full samples, but the differences are not statistically different from zero. Thus, this does not seem to be a problem in this paper. The results from these estimations are available from the authors upon request.

The point estimate of the interaction between the reform and $\ln markup$ in Table 2 shows that the effect of the reform was more pronounced if the markup was large. However, this result is not statistically significant at conventional levels. Neither does the indirect effect from equation (8) (as proxied by $diff \ln markup$ in our empirical model) show any significant impact on the reform-effect on pharmaceutical prices. Finally, expanding the informed section of the market through the introduction of the reform is found to increase the impact of the number of products on pharmaceutical prices, while the number of products had no significant effect on the prices themselves.

We estimate the basic model with and without n in order to study if the inclusion of this variable affects the estimate of the reform-effect. As can be seen in Table 2, the parameter estimate for the number of products is not statistically significant, and the estimated reform-effect is similar in size in both models. As such, the reform-effect does not seem to be affected by an exclusion of this variable. The estimated reform-effect in these two models is slightly lower than in the full model, but still approximately 2 percent. There is also a negative time trend of 0.2 percent per month.

The results from the model with the adjustment variable show that the estimated reform-effect is larger than in the basic model, with an estimated average price reduction during the study-period of 4 percent. One explanation for this result is that including the adjustment parameter in the empirical specification changes the parameter estimate for the time trend from -0.20 to -0.15 percent per month. The estimate for the reform dummy in this model, minus 6 percent, can be interpreted as an estimate of the long run effect of the reform. The relatively large standard errors of this estimate, nearly 2 percent, is likely caused by that the correlation between D and $D/(t-R)^\mu$ is as high as 0.82 (e.g. Greene, 2003, Chapter 4). According to a likelihood-ratio test, we can reject the hypothesis of an instant adjustment at the time of the reform. We therefore regard the estimates from the model with the adjustment variable as the more reliable ones. The estimated reform-effect for the model with adjustment is illustrated in Figure 1.

Figure 1 about here.

The results concerning generic pharmaceutical prices presented in Table 3 indicate an average price reduction due to the reform of one percent in all three estimated models not including the adjustment variable. Further the estimates show a negative price trend of 0.3 percent per month. As such, the reform-effect without adjustment is smaller in size for generics than for brand name pharmaceuticals, while the price trend is larger. Neither of the differences is statistically significant.

Table 3 about here.

The point estimates of the derivatives at the bottom of Table 3 show that both high values of *info* and $\ln markup$ enhance the effect of the reform as expected, but the effect of *info* is not statistically significant at conventional levels. For generics, the parameter estimate for $diff \ln markup * D$, is negative and statistically significant at the five percent level. Our interpretation of this is that a high markup prior to the reform of the competitors to generic *i* increases the reform-effect on the competitors, which in turns increases the need for generic *i* to lower its price. As discussed in Appendix 4, potential endogeneity will cause the estimators related to this variable to be positively biased. As such, this gives further support that the results are in line with our theoretical model. Finally, as in the estimations concerning brand name pharmaceuticals, expanding the informed section of the market through the introduction of the reform is found to increase the impact of the number of products on pharmaceutical prices, but the number of products had no significant effect on the prices themselves.

Turning to the results from our model including the adjustment process, the results show that the reform-effect is larger than in the basic models, with an estimated average price reduction of 4 percent. As for brand name pharmaceuticals, a large part of the difference between the models can be explained by the difference in the parameter estimates for the time trend. The point estimates for the parameter μ , indicate that the adjustment path for generics is less curved than that for brands. However the divergence is not statistically significant different from zero. Further, the low value of μ results in a correlation

between D and $D/(t - R)^\mu$ of 0.99, which increases the standard errors for the parameter estimates related to these variables (e.g. Greene, 2003, Chapter 4). The standard errors for the two estimates are over 5 percent and it is therefore not meaningful to discuss these estimates. However, the standard error for the total reform-effect is not influenced by the high correlation between D and $D/(t - R)^\mu$. As for the brand name sample, we can reject the hypothesis of an instant adjustment at the time of the reform, and we therefore regard the estimates from this model as the most reliable ones. The estimated reform-effect for the model with adjustment is illustrated in Figure 2.

Figure 2 about here.

Taken together, the results indicate that the effects of the reform on the price paths of pharmaceutical products are more similar for brand and generic products than expected, with an average price reduction due to the implementation of the reform of about 4 percent. These results could be compared to previous studies concerning the effects of substitution reforms on pharmaceutical prices (e.g. Buzzelli et al., 2006). In comparison, their results show an average price reduction of 3 percent after the implementation of the substitution reforms. It should, however, be noted that they use pharmaceutical price index data, from 16 OECD countries, and that their estimate therefore is not directly comparable to ours. Firstly, their estimate is an average for all drugs, not only those directly affected by the reforms. Secondly, the study by Buzzelli et al. does not incorporate any adjustment process in the empirical specification.

5 Discussion

The introduction of the substitution reform in 2002 changed the pharmaceuticals market in Sweden in three ways. First, due to the requirement that patients be informed if there were cheaper available products available, the amount of consumer information concerning pharmaceutical prices increased. Second, the out-of-pocket cost for patients not buying the cheapest generic increased somewhat (on average with approximately 12.5 SEK \approx 1.3 EURO) when the reference price was abolished. Third, the transaction costs for generic substitution

decreased when the reform was introduced. However, even before the reform a prescribed pharmaceutical product could be substituted for a cheaper generic version, although this required that the prescribing physician had given his/her consent to this on the prescription, or the patient requested substitution. Since the change in out-of-pocket cost was small and generic substitution could be done even before the reform, we argue that the major change due to the reform was the increase in consumer information about pharmaceutical prices.

As such, the introduction of the Swedish substitution reform was used as a natural experiment regarding the effects of increased consumer information on brand name and generic pharmaceutical prices. The main hypothesis to be tested was if the introduction of the substitution reform had led to any reduction in the price of brand name and/or generic pharmaceuticals. Another hypothesis to be tested was whether the possible price response due to the introduction of the reform differed between brand name and generic drugs.

First, following Frank and Salkever (1992), a theoretical model was set up to analyze the effects of increased consumer information on pharmaceutical prices. The model indicated that an increase in information would, under reasonable assumptions, led to a price reduction for brand name products, while the results for generic pharmaceuticals were more ambiguous. Based on our model, we also expected that the reduction in price would be large if the markup over marginal cost was large and/or the market had been characterized by low levels of information before the reform.

The results from the empirical part of the paper show an average reduction in prices due to the reform of about 4 percent during the study-period, both for brand name and generic pharmaceuticals. In addition, the results give some support for the reform-effect being amplified for pharmaceuticals in markets which had previously been characterized by low levels of consumer information, as well as for pharmaceuticals which prior to the reform had high markups over marginal cost.

We are also able to present tentative results regarding how an increase in information will affect the impact of the number of products on pharmaceutical prices. The results presented in this paper are in line with Grabowski and Vernon's (1992) conjecture, that expanding the informed portion of the market should increase the price lowering effect of entry.

A final point that needs to be discussed is how the brand name product

has previously been able to set a high price relative to the generic substitutes without losing market share in the absence of the substitution reform. Stiglitz (1979, p 340) suggests that “a flow of ignorance can be maintained either by entry of new firms or new individuals”. This could perhaps make it possible for, for example, a brand name producer to price its product high relative to its generic substitutes without losing market shares. To empirically investigate this question is outside the scope of the present paper, but warrants future research.

Appendix 1. Indirect effects

The effect on the price of generic i of a change in the price of the other generics is written

$$\frac{dP_i}{dP_j} = \left[\left(P_i - \frac{dC_i}{dQ_i} \right) \alpha \frac{d^2 D_i^I}{dP_i dP_j} + \alpha \frac{dD_i^I}{dP_j} \right] / (-\delta_i).$$

The first term includes $d^2 D_i^I / dP_i dP_j$, which describes how the slope of the demand curve facing firm i is affected by a change in the price of other generics. This term will be negative if a price reduction of other drugs changes the marginal consumer to one who is less price sensitive than the previous one. However, without knowing the specific form of the demand function, this term cannot be signed. Since the drugs are substitutes, the second term is positive and states that if the other prices are reduced, this will lower the demand for drug i , which works for a price reduction of drug i . The price response of firm i to a price change by the brand name firm and the price response of the brand name firm to a price change by the generic ones can be written correspondingly. Theoretically, we cannot rule out the possibility that the first term is negative and dominates the second one without making more explicit assumptions about the demand function. In this paper we will assume that $\frac{dP_i}{dP_j} > 0$, $\frac{dP_i}{dP_b} > 0$ and $\frac{dP_b}{dP_a} > 0$, without assuming a specific demand function. Among the possible demand functions that would result in the positive derivatives that we assume, a simple linear demand function like $D_i^I = d_1 - d_2 P_i + d_3 P_j + d_4 P_b$ can be mentioned.

Appendix 2. Second order conditions

The second order profit maximization condition for generic firm i , is written

$$\begin{aligned} d^2 \pi_i / dP_i^2 &= \delta_i = \left(P_i - \frac{dC_i}{dQ_i} \right) * \left[\alpha \frac{d^2 D_i^I}{dP_i^2} + (1 - \alpha) \frac{d^2 D_i^{UI}}{dP_i^2} \right] \\ &+ 2 * \left[\alpha \frac{dD_i^I}{dP_i} + (1 - \alpha) \frac{dD_i^{UI}}{dP_i} \right] < 0. \end{aligned}$$

The corresponding expression for the brand name firm is written

$$\begin{aligned}
 d^2\pi_b/dP_b^2 &= \delta_b = \left(P_b - \frac{dC_b}{dQ_b} \right) * \\
 &\left[\alpha \left(\frac{\partial^2 D_b^I}{\partial P_b^2} + 2 \frac{\partial^2 D_b^I}{\partial P_g^2} \left(\frac{\partial P_g}{\partial P_b} \right)^2 + \frac{\partial D_b^I}{\partial P_g} \frac{\partial^2 P_g}{\partial P_b^2} \right) + (1 - \alpha) \frac{dD_b^{UI}}{dP_b} \right] \\
 &+ 2 * \left[\alpha \left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g} \frac{\partial P_g}{\partial P_b} \right) + (1 - \alpha) \frac{dD_b^{UI}}{dP_b} \right] < 0.
 \end{aligned}$$

Appendix 3. Definitions of empirical measures

Denote the lowest observed price in exchange group g prior to the reform by $MinPrice_g$. Let $Sales_{it}$ denote the sales of pharmaceutical i at period t and let $Sales_{gt}$ be the corresponding for all pharmaceuticals in an exchange group. Then $info_g$, $\ln markup_i$ and $diff \ln markup_i$ is written as follows

$$\begin{aligned}
 info_g &= \left[\sum_{t=1}^R \left(\frac{\sum_{i=1}^n Price_{it} * W_{it}^1}{MinPrice_{gt}} \right) * W_{gt}^1 - 1 \right] * 100 \\
 \ln markup_i &= \ln \left[\sum_{t=1}^R Price_{it} * W_{it}^2 - MinPrice_g \right] \\
 diff \ln markup_i &= \ln \left[\sum_{i=1}^n \sum_{t=1}^R Price_{it} * W_{it}^3 - MinPrice_g \right] \\
 &\quad - \ln markup_i,
 \end{aligned}$$

where R is the time for the reform and

$$\begin{aligned}
 W_{it}^1 &= Sales_{it}/Sales_{gt}; & W_{gt}^1 &= Sales_{gt}/\left(\sum_{t=1}^R Sales_{gt} \right); \\
 W_{it}^2 &= Sales_{it}/\left(\sum_{t=1}^R Sales_{it} \right); & W_{it}^3 &= Sales_{it}/\left(\sum_{t=1}^R Sales_{gt} \right).
 \end{aligned}$$

That is, $info_g$ is defined as the extra expenditure on pharmaceuticals in exchange group g prior to the reform compared to the expenditure if the given quantity would have been bought at the lowest price in the exchange group each month, measured in percent. $\ln markup_i$ is measured as the logarithm of

the difference between the average price for pharmaceutical product i in the pre-reform period and the minimum price in the group of exchangeable products during that period. Finally, $diff \ln markup_i$ is created by subtracting $\ln markup_i$ from the logarithm of the difference between the average and the lowest price in one exchange group in the pre-reform period.

Appendix 4. Endogeneity

By recalling the definition of the variable $info$ from Appendix 3, we see that this variable depends on prices and therefore could be correlated with the error term of equation (12). For brand name pharmaceuticals, an increase in the pre-reform error terms will lead to an increase in $info$. The inclusion of product-specific fixed effects operates so that the sum of the error terms for a pharmaceutical during the whole study-period approaches zero. Therefore an increase in the pre-reform error terms for a pharmaceutical will lead to a reduction in the post-reform error terms for that pharmaceutical. For brand name firms, this results in a negative correlation between the post-reform error terms and the variable $info$, which will lead to negative bias in the estimators related to $info$. (The correlation between $info$ and the pre-reform error terms does not directly affect the bias, since $info$ is only included in the model together with the reform dummy.)

For the cheapest generic in an exchange group, an increase in the pre-reform error terms will lead to a reduction in $info$, since the denominator will increase proportionally more than the numerator. Therefore, the post-reform error term of this pharmaceutical will be positively correlated with $info$. This is also true for a uniform increase of the pre-reform error terms for all generics in an exchange group. In the estimation for generic pharmaceuticals, the estimators related to $info$ are therefore likely to be positively biased.

To sum up, for the reasons mentioned above we expect the estimators related to $info$ to be negatively biased for brands and positively biased for generics.

All pre-reform error terms for a pharmaceutical have a positive effect on the variable $\ln markup$, except the error term for the observation that affect $MinPrice$ which has a negative effect. Since in each exchange group there is only one observation of $MinPrice$, and several observations of other pre-reform prices, it is reasonable to expect the correlation between the lowest pre-reform

price and the post-reform error terms to be smaller in absolute size than the correlation between all pre-reform prices and these error terms. The variable $\ln markup$ is therefore expected to be negatively correlated with the post-reform error terms, making the estimators related to this variable negatively biased. Since an increase in $\ln markup$ will lead to a decrease in $diff \ln markup$, the estimators related to the latter variable will be positively biased.

The variable n is likely to be endogenous as well. An increase in the size of the error term will, through its effect on the price, lead to an increase in the number of products. This will, in turn, lead to positive bias in the estimators related to n . Including n in the estimation and treating it as exogenous might also lead to a negative bias in the estimator of the reform-effect. The reason is that the reform is likely to have a negative effect on prices, which in turn could have a negative effect on the number of products. If the number of products is estimated to have a negative effect on the prices, conditioning on it will then work for an overestimation of the negative price effect of the reform. However, in addition the reform makes it more likely that a generic product will be dispensed, increasing the probability of generic entry. Including n could therefore also lead to an underestimation of the negative price effect of the reform.

Appendix 5. Measuring the reform-effect

Denote the reform indicator variable by the vector \mathbf{D} . Let $\mathbf{X}(\mathbf{D})$ denote the $it \times 7$ matrix for the interaction variables ($info * D$), ($\ln markup * D$), ($\ln markup * info * D$), ($diff \ln markup * D$), ($diff \ln markup * info * D$), ($n * D$) and ($n * info * D$) and let $\boldsymbol{\gamma}$ be the vector of parameters related to these variables. Ignore for simplicity that our model also includes the variables n and $Trend$ and product-specific fixed effects, θ_i . Then our structural model is written

$$\ln \mathbf{Price} = \mathbf{X}(\mathbf{D})\boldsymbol{\gamma} + \mathbf{D}\beta_1 + \boldsymbol{\varepsilon}.$$

If this equation is estimated on reduced form, excluding $\mathbf{X}(\mathbf{D})$, the estimator for β_1 becomes according to Greene (2003, Chapter 8).

$$b_1 = \beta_1 + (\mathbf{D}'\mathbf{D})^{-1}\mathbf{D}'\mathbf{X}\boldsymbol{\gamma} + (\mathbf{D}'\mathbf{D})^{-1}\mathbf{D}'\boldsymbol{\varepsilon}.$$

Now, in order to simplify things assume that there are observations in only two periods, one before the reform ($t = 1$) and one after ($t = 2$), and let us

calculate what b_1 represents if we omit the interaction terms from our empirical specification. The matrices are written

$$\mathbf{D}' = \begin{pmatrix} \mathbf{0} & \mathbf{1} \end{pmatrix}; \mathbf{D} = \begin{pmatrix} \mathbf{0} \\ \mathbf{1} \end{pmatrix}; \mathbf{X} = \begin{pmatrix} \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{x}_1 & \mathbf{x}_2 & \mathbf{x}_3 & \mathbf{x}_4 & \mathbf{x}_5 & \mathbf{x}_6 & \mathbf{x}_7 & \mathbf{x}_7 \end{pmatrix}; \boldsymbol{\varepsilon} = \begin{pmatrix} \boldsymbol{\varepsilon}_1 \\ \boldsymbol{\varepsilon}_2 \end{pmatrix},$$

where $\mathbf{0}$ and $\boldsymbol{\varepsilon}_1$ are vectors of zeros and error terms for the I pharmaceuticals before the reform, and where $\mathbf{1}$, \mathbf{x}_1 , $\mathbf{x}_2 \dots \mathbf{x}_7$ and $\boldsymbol{\varepsilon}_2$ are vectors of ones, the interaction variables and the error terms for the I pharmaceuticals after the reform. Matrix multiplication then gives

$$b_1 = \beta_1 + \sum_{k=1}^7 \frac{1}{I} \sum_{i=1}^I x_{ki} \gamma_k + \frac{1}{I} \sum_{i=1}^I \varepsilon_{i2}.$$

That is, if we exclude the interaction terms with the reform dummy, the average effect of these variables will still be captured by the parameter estimate of the reform indicator variable.

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Table 1. Descriptive statistics

Variable	Brands		Generics	
	Mean	Std. Dev.	Mean	Std. Dev.
<i>Price</i>	1796.04	4359.74	1727.62	5137.14
$\ln Price$	6.64	1.18	6.39	1.31
<i>D</i>	0.67	0.47	0.67	0.47
<i>info</i>	7.98	22.51	12.90	81.63
$\ln markup$	4.34	1.49	3.21	1.62
$diff \ln markup$	-0.28	0.59	0.97	1.12
<i>n</i>	2.35	1.44	3.29	1.56
<i>Trend</i>	33.99	20.05	33.09	19.45
$info * D$	5.15	16.44	8.74	65.17
$\ln markup * D$	2.90	2.37	2.10	1.99
$\ln markup * info * D$	26.46	101.03	32.65	371.04
$diff \ln markup * D$	-0.18	0.51	0.69	1.08
$diff \ln markup * info * D$	-1.10	9.34	13.91	99.60
$n * D$	1.66	1.70	2.24	2.07
$n * info * D$	15.02	55.61	28.86	119.17
Nr. $\ln markup$	47008		57861	
Nr.	47121		58466	
Products	873		1613	

Nr. $\ln markup$ is the number of observation for the variables including $\ln markup$ or $diff \ln markup$ while Nr. is the number of observations for all other variables.

Table 2. Brand name prices

	Full	Basic+n	Basic	Basic+adj.
<i>D</i>	22.32** (11.13)	-18.85*** (5.71)	-18.86*** (5.70)	-56.88*** (17.19)
<i>Trend</i>	-2.01*** (0.51)	-2.01*** (0.52)	-2.01*** (0.50)	-1.53*** (0.47)
<i>n</i>	1.60 (1.36)	-0.23 (0.79)		
<i>info * D</i>	-4.35* (2.48)			
<i>ln markup * D</i>	-6.25** (2.94)			
<i>ln markup * info * D</i>	0.71* (0.39)			
<i>diff ln markup * D</i>	1.98 (5.84)			
<i>diff ln markup * info * D</i>	1.20* (0.67)			
<i>n * D</i>	-2.28 (1.68)			
<i>n * info * D</i>	-0.04 (0.05)			
<i>D/(t - R)</i> (γ)				38.14*** (13.89)
<i>D/(t - R)</i> (μ)				354.00*** (see note)
$\Delta \ln Price / \Delta D$	-19.25*** (5.60)	-18.85*** (5.71)	-18.86*** (5.70)	-42.40*** (12.25)
$d \ln Price / d(info * D)$	-1.73* (0.99)			
$d \ln Price / d(\ln markup * D)$	-0.62 (4.12)			
$d \ln Price / d(diff \ln markup * D)$	11.59 (7.37)			
$d \ln Price / d(n * D)$	-2.62* (1.54)			
$d \ln Price / dn$	-0.15 (0.77)	-0.23 (0.79)		
Sample size	47008	47121	47121	47121
Log likelihood	76724	76858	76859	76868

The reported values are the estimates multiplied by 1000. The regressions include product-specific fixed effects. Robust standard errors, conditioned on the parameter, μ , are reported in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. The 95%-confidence interval for μ is $0 < \mu < 1086$. The differential/derivatives are evaluated at the mean for each variable. For $D/(t-R)^\mu$ the mean in the post-reform period is used.

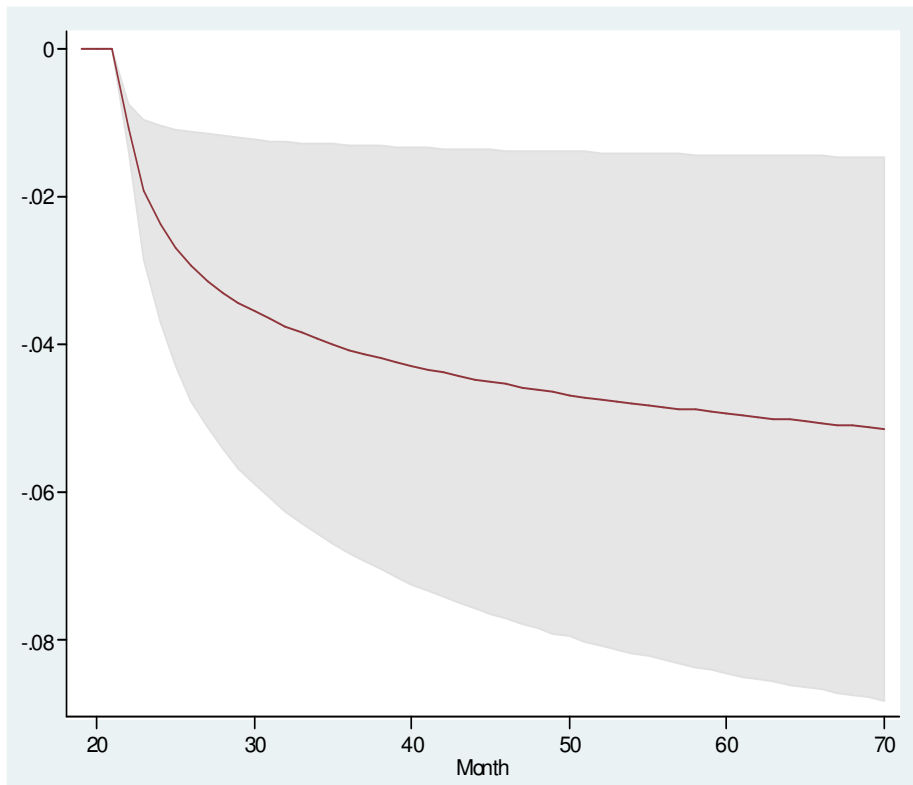


Figure 1. Brand name prices, estimated reform effect with 95 percent CI

Table 3. Generic prices

	Full	Basic+n	Basic	Basic+adj.
<i>D</i>	35.85*** (12.08)	-11.02*** (1.70)	-10.98*** (1.70)	-135.85** (55.49)
<i>Trend</i>	-2.67*** (0.75)	-2.75*** (0.76)	-2.74*** (0.76)	-2.01*** (0.62)
<i>n</i>	2.39*** (0.73)	0.75 (0.57)		
<i>info * D</i>	-1.37 (1.03)			
<i>ln markup * D</i>	-8.03*** (2.38)			
<i>ln markup * info * D</i>	0.22 (0.17)			
<i>diff ln markup * D</i>	-10.82*** (4.16)			
<i>diff ln markup * info * D</i>	0.19 (0.14)			
<i>n * D</i>	-2.50** (1.05)			
<i>n * info * D</i>	0.01 (0.01)			
<i>D/(t - R)</i> (γ)				125.20** (54.51)
<i>D/(t - R)</i> (μ)				101.00*** (see note)
$\Delta \ln Price / \Delta D$	-11.19*** (1.79)	-11.02*** (1.70)	-10.98*** (1.70)	-41.34*** (14.38)
$d \ln Price / d(info * D)$	-0.44 (0.36)			
$d \ln Price / d(\ln markup * D)$	-5.13** (2.53)			
$d \ln Price / d(diff \ln markup * D)$	-8.40** (4.09)			
$d \ln Price / d(n * D)$	-2.41** (1.01)			
$d \ln Price / dn$	0.77 (0.58)	0.75 (0.57)		
Sample size	57861	58466	58466	58466
Log likelihood	79506	79847	79843	79855

The reported values are the estimates multiplied by 1000. The regressions include product-specific fixed effects. Robust standard errors, conditioned on the parameter, μ , are reported in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. The 95%-confidence interval for μ is $0 < \mu < 672$. The differential/derivatives are evaluated at the mean for each variable. For $D/(t-R)^\mu$ the mean in the post-reform period is used.

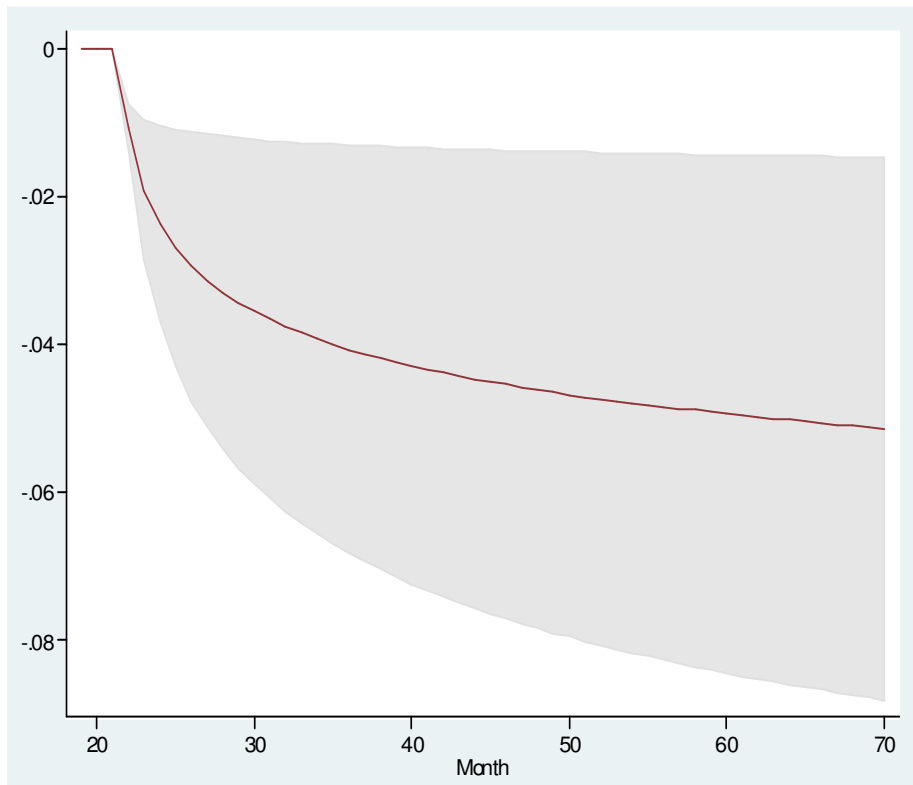


Figure 2. Generic prices, estimated reform effect with 95 percent CI