DEFINING RELEVANT MARKETS FOR PHARMACEUTICALS

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ABSTRACT

To identify the relevant product markets for Swedish pharmaceuticals, a spatial econometrics approach is employed. First, we calculate Moran’s Is for different market definitions and then we use a spatial Durbin model to determine the effect of price changes on quantity sold of own and competing products. As expected, the results show that competition is strongest between close substitutes; however, the relevant product markets for Swedish pharmaceuticals extend beyond close substitutes down to products included in the same class on the four-digit level of the Anatomic Therapeutic Chemical system as defined by the World Health Organization. The spatial regression model further indicates that increases in the price of a product significantly lower quantity sold of that product and in the same time increase the quantity sold of competing products. For close substitutes (products belonging to the same class on the seven-digit level of the Anatomic Therapeutic Chemical system), as well as for products that, without being close substitutes, belong to the same therapeutic/pharmacological/chemical subgroup (the same class on the five-digit level of the Anatomic Therapeutic Chemical system), increased competition is also visible after 1 July 2009 when the latest policy changes with regards to pharmaceuticals have been implemented in Sweden.

Keywords: competition, horizontal merger guidelines, market delineation, market reform, spatial Durbin model
JEL classification numbers: L11, L25, L26, L81

I. INTRODUCTION

Several previous papers have studied competition between brand name and generic pharmaceutical products, with focus on the relevance of consumer information (Frank and Salkever, 1992; Grabowski and Vernon, 1992; Sorensen, 2000; Granlund and Rudholm, 2011), the importance of the physician in the choice of pharmaceuticals (Hellerstein, 1998; Coscelli, 2000; Richard and Van Horn, 2004; Granlund and Rudholm, 2012), or the relationship between brand name prices and generic market shares (Aronsson et al., 2001; Rizzo and Zeckhauser, 2005). Yet other papers have focused on competition between different brand name pharmaceuticals.
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within one or a few specific fields of treatment (Berndt et al., 1995; Ellison et al., 1997; Rudholm, 2003; Richard and Van Horn, 2004).

Frank and Salkever (1992) showed that an increase in the number of consumers informed about the availability of cheaper generic alternatives would 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 generic entry on brand name pharmaceutical prices, Frank and Salkever showed that, theoretically, this was not necessarily the case. 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 (i.e., brands and generics) were found to be large. In fact, the highest listed price exceeded on average the cheapest available alternative by as much as 50 percent. In addition, the results gave 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. Finally, Granlund and Rudholm (2011) studied the impact of increased consumer information on the price of brand name and generic products. In a theoretical model, they showed that an increase in information would, under reasonable assumptions, lead to a price reduction for brand name products, while the results for generic pharmaceuticals were more ambiguous. The results from the empirical part of their paper showed an average reduction in prices due to the an information increasing reform in the Swedish pharmaceuticals market of about four percent during the study period, both for brand name and generic pharmaceuticals. In addition, their results gave 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.

Another topic under study has been the importance of the physician in the brand name or generic prescription decision (Hellerstein, 1998; Coscelli, 2000; Richard and Van Horn, 2004; Granlund and Rudholm, 2012). Using data from a survey of physicians and their patients, Hellerstein (1998) found significant differences between physicians’ likelihood to prescribe generics and also found that it was difficult to determine why some physicians were more likely to prescribe generic drugs. Coscelli (2000) used information about doctor and patient characteristics, as well as information about when and how patients switched physicians, to estimate the probability of a switch of pharmaceutical brands. Her results showed that there was persistence in the use of pharmaceuticals for both patients and physicians. Richard and Van Horn (2004) also found that there was persistence in physician prescription choices, especially for incumbent products. They suggested that this was caused by incumbent products having a larger installed base of patients rather than advertising generating goodwill for such products. Finally, Granlund and Rudholm (2012) studied the prescribing physician’s influence on consumer choice between medically equivalent pharmaceuticals. In their paper, if consumers cared about the advice of their prescribing physician, they were expected to allow substitution to cheaper products most often for generics since the names of these products were so similar that consumers might have considered them to be the same pharmaceutical. However, when consumers were offered a cheaper substitute instead of a brand or secondary brand with a distinct product name, consumers might have believed that the substitute was different from the pharmaceutical that their physician had prescribed because of the difference in names. As expected from the above discussion, the results from the empirical part of Granlund and Rudholm (2012) also showed that the predicted probability of opposing substitution was considerably higher for brands and secondary brands than for generics. For the whole population, the predicted probability of opposing substitution was 10 percent for a generic and 24–25 percent for a brand or a secondary brand.
A third topic of study has been the relationship between brand name prices and generic market shares. Aronsson et al. (2001) set up a model where the relative change of market share of the original drug depended on the price of the original relative to the price of the generic substitutes. Their empirical results showed that a higher price of the original product, relative to the average price of the generic substitutes, significantly decreased the market share for the original product for five out of 12 studied substances. Also, the results indicated that the introduction of a so-called reference price system in 1993 lowered pharmaceutical prices in Sweden. Rizzo and Zeckhauser (2005) showed that the higher the share of prescriptions that was filled by generics, the lower the average brand-name price to consumers. They suggested that this was due to consumers’ becoming more likely to substitute brand-name drugs for generics when the price gap was great. They found that this effect was large, with a 10 percent increase in generic script share associated with a 15.6 percent decline in the average price paid for brand-name drugs.

All papers mentioned above explicitly or implicitly assume that the relevant market to study is where a brand name product is in competition with one or more generic products. Only a few previous studies of the pharmaceuticals market analyse competition between different pharmaceutical substances (e.g., Berndt et al., 1995; Ellison et al., 1997; Rudholm, 2003; Richard and Van Horn, 2004). The general conclusion from these papers is that there is competition also between substances, but that the degree of competition is considerably lower than between brands and generics for the same active substance.

It should, however, be noted that all papers discussed above, including the studies of competition between substances, set out by assuming a relevant product market to be studied, instead of first trying to empirically decide the relevant scope of the product market. The purpose of our paper is thus to use spatial econometric methods as a way to decide the scope of relevant product markets for pharmaceuticals in Sweden. This is, to our knowledge, the first attempt at defining the relevant product market for pharmaceuticals in Sweden.

We set out by considering the basic definition of the area of a market: a market for a good is the area within which the price of a good tends to uniformity, allowance being made for transportation and transaction costs. Under such circumstances, price movements will be highly correlated. If one seller changes the price of the product, other sellers will have to follow suit, and the cross-price elasticities of demand will be positive and statistically significant. This should also be true in product space; if two imperfect substitutes are similar enough that a price change of one product forces the sellers of the substitute product to change their prices as well, we will consider the products to be competitors in the same product market.

Our results show that price movements are highly positively correlated when the market is defined as brands and generics for a certain pharmaceutical substance (seven-digit level in World Health Organization’s Anatomic Therapeutic Chemical (ATC) classification system), indicating that such products are competing in the same product market. However, the results also show clearly correlated price movements in broader market definitions (five- and four-digit ATC levels), indicating that the relevant product market for pharmaceuticals in Sweden extends beyond the seven-digit ATC level containing brands and generics of the same pharmaceutical substance.

In this paper, we also investigate the effects of the policy changes from 2009 and 2010 in terms of possible changes in the price elasticity of demand for pharmaceutical products. The results show a pronounced increase in competition (measured as the cross-price elasticity of demand) between brands and generics for a certain pharmaceutical substance (seven-digit ATC level) after 1 July 2009, a slight increase in competition at the five-digit ATC level, and no significant change at the four-digit ATC level.

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1 This definition dates back to Cournot, and is discussed more in detail in Stigler and Sherwin (1985).
The Swedish pharmaceuticals market is presented in the next section. Section III describes the statistical method, while the data employed and descriptive statistics are presented in section IV. The results from our study are presented in section V. Finally, section VI summarizes the findings and draws conclusions.

II. THE SWEDISH PHARMACEUTICALS MARKET 1993–2013

Before 2002, a reference price system introduced in January 1993 was in effect for pharmaceuticals in Sweden. 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, p. 20, 1996, p. 31). The effects of the reference price system on pharmaceutical prices have been previously analysed (see e.g., Aronsson et al., 2001 and Bergman and Rudholm, 2003). The results from Aronsson et al. showed that the introduction of the reference price system reduced brand name pharmaceutical prices where generic substitutes were available at the introduction of the system. However, Bergman and Rudholm found that the Swedish reference price system had been effective only for those products which already faced generic competition at the introduction of the system. Before the reference price system was introduced, generic entrants typically set prices far below the brand name product, without capturing more than a relatively small fraction of the market (Aronsson et al., 2001). When the system was introduced, the brand name manufacturers were forced to lower their prices to a level close to that of the generics, in order for consumers to be fully reimbursed.

The Swedish pharmaceuticals market was reformed on 1 October, 2002. The reforms require that from that date pharmacists inform the consumers if there are substitute products available, as well as that the cheapest available generic substitute or parallel imported product (which is considered to be perfect substitute for the brand name drug by the Swedish Medical Products Agency) will be provided within the Swedish pharmaceuticals insurance system. During the period 2002 and until 2009, ‘available’ was interpreted as that the cheapest available product at the local pharmacy should be dispensed to the consumers. The pharmacist also had to inform the consumers that they could buy the prescribed pharmaceutical product instead of the suggested generic if they paid the difference in price between the products themselves. Finally, the rules that have been in effect since 2002 require that pharmacists substitute the prescribed pharmaceutical product to the cheapest available generic or parallel imported product in all 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 alternative. In cases where the physician prohibits the switch due to medical reasons the consumer is still reimbursed by the insurance system.

It should be noted that even before the 2002 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 the fact. Since these measures were abolished in 2002, this means that the transaction cost of generic substitution was lowered somewhat, which could also affect the pricing behaviour of pharmaceutical firms.

Both before and after the substitution reform of 2002, pharmaceutical firms in Sweden have had the formal right to decide the price of their products, but for a pharmaceutical to be included in the Swedish pharmaceuticals insurance system, its price had to be authorized. The prices of
new pharmaceutical products are authorized using a cost-benefit analysis that has to be provided with the application for the right to sell the pharmaceutical product within the Swedish insurance system. During the period 1969 until 2010, pharmaceuticals were sold through a nationwide government owned monopoly, the National Corporation of Swedish Pharmacies (NCSP). The NCSP had a margin on the pharmaceutical products that they sold that was determined by the Pharmaceutical Benefits Board. The regulations also implied that the NCSP was required to charge a nationwide uniform price for each pharmaceutical product in Sweden. Thus, the relevant geographic market for pharmaceutical products is, by regulation, Sweden as a whole.

The 2002 reform also affected the out-of-pocket cost for patients as the rules regulating this were changed when the reference price was abolished. Under the substitution reform costs up to 100 percent of the cheapest generic alternative are included in the pharmaceutical insurance system, compared to 110 percent during the reference price system. This increased 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 (≈ 2 EURO, exchange rate 2013-06-28) per pharmaceutical (Granlund and Rudholm, 2011).

In 2009 and 2010, the Swedish pharmaceuticals market was reformed again, and this time more substantially than in 2002. From 1 July, 2009, the maximum price of pharmaceutical products that had previously been sold under patent protection was set at 35 percent of the price 12 months before patent expiration. According to the Swedish Dental and Pharmaceutical Benefits Agency (DPBA), this resulted in a decrease in pharmaceutical costs of approximately 350 million SEK (≈ 50 million USD, exchange rate 28 June, 2013). On 1 November, 2009, the trade margin on pharmaceutical products was increased by a fixed amount of 10 SEK per prescription of a generic product or a brand name product that was exposed to generic competition. This change increased the costs of the trade margin with approximately 460 million SEK (≈ 64 million USD, exchange rate 28 June, 2013). The purpose of this change was twofold. The decision makers wanted to promote generic substitution and also increase the compensation for the pharmacies in order to increase the number of pharmacies in Sweden. From October 2009, the interpretation of cheapest available generic was changed from the cheapest product at the local pharmacy to the cheapest product in the market as a whole. The cheapest product, called ‘product of the month’ was determined by the DPBA in an auction where the lowest bid wins. Then, 3 May, 2010, the number of products labelled ‘product of the month’ has been increased to three, but the pharmacies are obliged to follow the DPBA ranking of the products and should only switch to one of the other two products in exceptional cases. Finally, in February 2010, the nationwide government owned monopoly that had been in effect since 1969 was abolished, and the NCSP was exposed to competition from other pharmacy chains. However, it should be noted that since prices on prescription pharmaceuticals were still regulated, competition between pharmacies was mainly in the pricing of OTC-drugs, but also in opening hours, quality of service, etc.

III. STATISTICAL METHOD

III.1 Spatial dependence

In this paper, we use spatial econometrics to determine the extent of the market for pharmaceutical products in Sweden and to identify the impact of the new regulations introduced during 2009 and 2010. Spatial econometrics has traditionally been used in real estate-related research to model property prices and to identify the impact of amenities and/or disamenities on own and neighbouring property values (Cohen and Coughlin, 2008; Conway et al., 2010;
The fundamental concept of spatial econometrics is that of spatial dependence. In real estate-related research, spatial dependence implies that the value of each property is influenced not only by its own structural, neighbourhood, and locational characteristics (e.g., the square footage of the land, the median household income in the neighbourhood, or the distance to the city centre), but also by the value of its neighbouring properties. Let us consider one of the most frequently used models in spatial econometrics, the spatial autoregressive model (SAR):

\[ Y = \rho WY + X\beta + \epsilon, \quad \epsilon \sim N(0, \sigma^2 I_n), \]  

where:

- \( Y \) is the dependent variable;
- \( WY \) is the spatial lag of the dependent variable;
- \( \rho \) is the spatial lag coefficient for the dependent variable;
- \( X \) represents a vector of the explanatory variables of the model;
- \( \beta \) represents a vector of coefficient estimates, indicating the strength and direction of the impact of the explanatory variables on \( Y \);
- \( \epsilon \) the error term, \( \epsilon \sim N(0, \sigma^2 I_n) \); and
- \( n \) is the number of observations.

Equation (1) shows that the data is simultaneously generated, with the value of \( Y \) depending on the value of its neighbours (through the spatial lag that includes a spatial weights matrix, \( W \)), or, in other words, that the analysed phenomenon is impacted by spatial dependence, or spatial externalities.

Disregarding this phenomenon, researchers may use Ordinary Least Squares (OLS) to estimate regression parameters in their studies. One of the assumptions of traditional OLS modelling is that of independence of observations, so that:

\[ E(\epsilon_i \epsilon_j) = E(\epsilon_i)E(\epsilon_j) = 0. \]  

This assumption does not hold in the presence of spatial externalities; consequently, the use of OLS produces biased and/or inconsistent coefficients estimates (Anselin, 1988; LeSage, 2008; Brown et al., 2009). To address this issue, OLS is replaced by a spatial regression model.

Although traditionally used for researching spatially-defined units such as real estate properties, spatial econometrics has also been employed as a method for analysing relationships between non-spatial units. Katz (1953) and Bonacich (1987) measure the degree to which an individual’s status in a network is a function of those to whom he or she is connected. They use a weight matrix to describe network connections and construct a status index based on the status of the first- and higher-order ‘neighbours’ in the network. The influence of higher-order ‘neighbours’ is weighted down to allow for the lower effectiveness of longer chains. Black (1992) also investigates network autocorrelation: the dependence of variable values on given links to variable values on other links in the same network. He differentiates this from spatial autocorrelation: variable values at given locations being influenced by variable values at neighbouring locations in a spatial context. He discusses the level of dependence in the residuals of an OLS model and corrects it by integrating an autoregressive term and, in a second model, by specifying the error term so as to involve a dependency structure. He then applies this method to study migration flows in the US using a non-spatial network context. Sonis et al. (2001) analyse feedback loops in inter-regional trade in an attempt to identify within-nation trading blocks. They determine the economic interdependencies in inter-regional trade at different levels of aggregation and create functional maps of the hierarchy of trade feedback loops to illustrate regions of intense trade. Ballester et al. (2006) study Nash equilibriums in social networks using the spatial autocorrelation concept. They relate individual outcome to the position in the network of local interactions and discuss the aggregate equilibrium of the network and what happens when this equilibrium is disrupted. Blankmeyer et al. (2007) use nearest-neighbours-based spatial regression modelling to study salary benchmarking in the US. Salary benchmarking is a practice that violates the assumption of independence in that
it adjusts top-level management compensation to reflect salaries offered by peer institutions (those against which firms compete when hiring top-level management). Instead of using the spatial location of firms, Blankmeyer et al. define neighbours on the basis of groups of peer institutions (based on the size of the firms, production processes, and resource or product markets).²

To our knowledge, no study has previously used spatial modelling to determine the relevant product market for pharmaceuticals or to investigate changes in competition between pharmaceutical products due to regulatory changes. Our study fills in this gap and the first step in this process is to define the neighbours or, in our case, the competing products, and based on this definition to create a spatial weights matrix to be used in the spatial model. Our data are defined bi-weekly over 132 cross-sections from 1 January, 2006 until 30 June, 2011. The frequent occurrence of product entry and exit in each cross-section makes the use of a spatial model for the whole panel difficult; instead, we investigate results for each cross-section in part and graph the results against time to emphasize trends in price competition and the impact of price on quantity sold of pharmaceutical products on the Swedish market.

III.2 The spatial weight matrix

The spatial weight matrix is used to define neighbours or, in our study, products that are competing and that are consequently likely to have an impact on each other’s prices and/or quantities sold. Pharmaceuticals commercialized on the Swedish market are classified according to the Anatomical Therapeutic Chemical classification system, and this will be used to create different spatial weight matrices.

The ATC system divides pharmaceuticals into categories according to their targeted organ/system and their therapeutic, pharmacological, and chemical properties. At the first level, the least detailed, drugs are divided into 14 main groups solely on the basis of their targeted organ/system. At the second, third, and fourth level, drugs are divided according to their therapeutic/pharmacological/chemical properties. At the fifth level, the most detailed, drugs are categorized by their active chemical substance (WHOCC, 2012). For example, the code N02BA01 corresponds to acetylsalicylic acid, commonly known as aspirin. According to the ATC system, this code shows that acetylsalicylic acid impacts the nervous system (N), is an analgesic (02), belongs to ‘other analgesics and antipyretics’ subgroup (B) and to its subdivision ‘salicylic acid and derivatives’ (A), and that its active chemical component is acetylsalicylic acid (01):

N Nervous system — 1st level: anatomical main group
   02 Analgesics — 2nd level: therapeutic subgroup
      B Other analgesics and antipyretics— 3rd level: pharmacological subgroup
         A Salicylic acid and derivatives— 4th level: chemical subgroup
            01 Acetylsalicylic acid — 5th level: chemical substance

We have created weight matrices for each bi-weekly cross-section and ATC level using MATrix LABoratory (MATLAB) so that the relationship between two products is defined as 1 if they are competing products and 0 if they are not. Two products are defined as competing when belonging to the same ATC group. Assuming that we have five products competing to each

²A description of studies focused on non-spatial structured dependence can be found in LeSage, 2008.
other as it follows: \( P_1 \) with \( P_2 \) and with \( P_3 \), and \( P_4 \) with \( P_5 \), the first-order contiguity weight matrix would be created as:

\[
\begin{pmatrix}
P_1 & P_2 & P_3 & P_4 & P_5 \\
P_1 & 0 & 1 & 1 & 0 \\
P_2 & 1 & 0 & 1 & 0 \\
P_3 & 1 & 1 & 0 & 0 \\
P_4 & 0 & 0 & 0 & 1 \\
P_5 & 0 & 0 & 0 & 1 & 0 \\
\end{pmatrix}
\] (3)

The spatial matrix \( M \) is symmetrical (if \( P_1 \) competes with \( P_2 \), then \( P_2 \) competes with \( P_1 \)) and has zero values on its diagonal, since a product cannot compete with itself. The matrix \( M \) can be standardized to have row sums of unity, generating the row-standardized spatial weights matrix \( W \):

\[
\begin{pmatrix}
P_1 & P_2 & P_3 & P_4 & P_5 \\
P_1 & 0 & 1/2 & 1/2 & 0 & 0 \\
P_2 & 1/2 & 0 & 1/2 & 0 & 0 \\
P_3 & 1/2 & 1/2 & 0 & 0 & 0 \\
P_4 & 0 & 0 & 0 & 1 & 0 \\
P_5 & 0 & 0 & 0 & 1 & 0 \\
\end{pmatrix}
\] (4)

The row-stochastic spatial weights matrix \( W \) is then multiplied by a \( 5 \times 1 \) vector \( Y \) representing the dependent variable to produce a spatial lag of the dependent variable taking the vector form \( WY \), which is a linear combination of the values of the dependent variable \( Y \) from the neighbouring observations. The spatial lag is then used in the spatial regression model. In our study, \( Y \) represents a vector of price or quantity sold of each pharmaceutical product:

\[
\begin{pmatrix}
P_1 & P_2 & P_3 & P_4 & P_5 \\
P_1 & y_1 \\
P_2 & y_2 \\
P_3 & y_3 \\
P_4 & y_4 \\
P_5 & y_5 \\
\end{pmatrix}
\] (5)

By standardizing the spatial weight matrix we compare the value of the spatial lag coefficient \( \rho \) to 1 when reading the regression results, which is easy to interpret. The spatial lag coefficient shows the strength of spatial dependence and is estimated in the spatial regression model with the aid of the spatial weights matrix.

As explained above, we define neighbours (competing products) according to the level of analysis. At the most detailed level of analysis we define drugs with the same seven-digit ATC code (e.g., all drugs coded as N02BA01) as neighbours, or, in other words, as competing. Consequently, the corresponding element in the matrix \( M \) (at the intersection of any row and column corresponding to two competing products) is 1. When moving up one step we define drugs with the same five-digit ATC code (e.g., all drugs coded as N02BA) as competing, and so on.

III.3 The Moran’s I statistic

To determine the level and significance of competition between pharmaceutical products we first use the Moran’s I statistic. This is the most common descriptive measure of spatial dependence and thus an indicator of the relationship between each item and its neighbours for
a given characteristic. In our case, the Moran’s I statistic is firstly used to determine whether changes in prices of pharmaceutical products are interrelated within the various pharmaceutical groups; secondly, the Moran’s I is used as an indicator of the magnitude and significance of the price relationship between competing products (and later of the relationship between quantities sold of pharmaceutical products). The Moran’s I is given by the expression:

\[ I = \frac{N}{\sum_i \sum_j w_{ij}} \times \frac{\sum_i \sum_j w_{ij} \times (X_i - \bar{X}) (X_j - \bar{X})}{\sum_i (X_i - \bar{X})^2}, \]  

(6)

where \( N \) is the number of analyzed units; \( X \) is the variable of interest (e.g., unit price); \( \bar{X} \) is the mean of the variable \( X \); \( w_{ij} \) are elements of the spatial weights matrix \( W \); and \( i = 1, \ldots, n \) and \( j = 1, \ldots, n \), where \( n \) is the number of observations.

Following Moran’s own definition from 1948, Dray et al. (2006) and Dray (2011) re-write the expression (6) for calculating the Moran’s I coefficient in matrix format and show that, since the matrix \( W \) can be any non-negative spatial weights matrix, the Moran’s I can be either positive or negative. For binary row-normalized weights matrices, like the one we use in the present study, the value of the Moran’s I statistic falls in between -1 and 1. A positive value indicates positive spatial autocorrelation. If price is the variable of interest and we consider the fact that we calculated Moran’s I values on price differences between two successive time periods, a high positive value of the Moran’s I statistic indicates that if the price of a product increases between two successive time periods, it is likely that the price of competing products also increases. The opposite is also true: if the price of a product decreases, it is likely that the price of competing products also decreases. A negative value of the Moran’s I statistic indicates negative spatial autocorrelation. In other words, if a price of a product increases, the price of competing products decreases and the other way around. If Moran’s I value is equal to -1, a map of the value of interest (e.g., price changes) for the analysed units (e.g., pharmaceutical products) would look like a check board, with black representing increase in value and white representing decrease in value. A statistically significant Moran’s I indicates significant autocorrelation between the analysed units. If the value of Moran’s I is not statistically significant we can conclude that there is no autocorrelation between the analysed items at the significance level chosen for analysis.

In our study the Moran’s I is used as a first indicator for the extent of the market. We investigate whether there is price competition between pharmaceutical products at each ATC level by calculating the corresponding Moran’s I. If we find that the Moran’s I for price changes has a high positive and significant value this means that products with that same ATC code are competing in the sense that if one product lowers its price the other products follow suit, and thus belong to the same market. We start by investigating whether there is significant price competition between products with the same seven-digit ATC code, as for example, products coded N02BA01 ‘acetylsalicylic acid’: Albyl minor, Aspirin, Bamyl, and Magnecyl brus (WHOCC, 2012). We expect strong competition at this level, since many of these medicines are close substitutes, and since previous literature have in most cases found strong competition at this level. We then move one level up to investigate competition between drugs which have the same first five digits of their ATC code. In other words, at this level we analyse price competition between drugs that belong to the same chemical subgroup, as for example drugs in the category N02BA ‘salicylic acid and derivatives’. The third step is to investigate price competition between drugs with the same first four digits of their ATC code, for example N02 ‘analgesics’. The last step is to consider products with the same first three digits of their ATC code, for example N02 ‘analgesics’. The last step is to investigate competition between drugs pertaining to the main therapeutic groups (same first digit of their ATC code). We expect that at one point between the first and the last
level in the analysis, competition between pharmaceuticals weakens sufficiently so that we can define the boundaries of the market for pharmaceuticals in Sweden.

The extent of the market is thus defined by that level where the Moran’s I statistic becomes either so small that it does not have any economic significance or when it becomes insignificant at the 95 percent confidence level. Since the Moran’s I is a type of correlation coefficient measuring the strength of the relationship between each observation and its neighbours, we argue, in accordance to Connolly (2007), that a Moran’s I value lower than 0.3 denotes weak product space dependence, a value between 0.3 and 0.6 denotes moderate product space dependence, and a value higher than 0.6 denotes strong product space dependence. Consequently, we imply that a value of Moran’s I that is insignificant or lower than 0.3 indicates very weak or no price competition. Thus, the market extent is defined by values of Moran’s I that are both statistically significant and higher than 0.3.

III.4 The spatial Durbin model

To analyse the impacts of the new regulations introduced in 2009 and 2010 regarding the sale of pharmaceuticals on the Swedish market we have chosen to use a spatial Durbin model (SDM). The need for the use of a SDM arises from the likely presence of omitted variables that may be correlated with the included variables. LeSage and Pace (2009) show that the presence of an omitted variable that exhibits spatial dependence and is correlated with one of the included explanatory variables requires the use of a spatial model that includes both a spatial lag of the dependent variable and of the explanatory variables (see also Kirby and LeSage, 2009). Economic theory also shows that, while the impact of a change in price on quantity sold of a product is usually negative, the impact of the same change in price on quantity sold of other products can be either positive or negative depending on if the products are substitutes or complements. These potentially complex relationships could not be modelled with the use of, e.g., a spatial autoregressive model (SAR), because in a SAR model the entire spatial dependence structure is reduced to one single coefficient (Vega and Elhorst, 2013); furthermore, the ratio between the indirect and direct effects is the same in a SAR model for every explanatory variable, which is unlikely in many empirical studies (LeSage and Pace, 2009; Elhorst, 2012; Pace and Zhu, 2012; Vega and Elhorst, 2013). LeSage and Pace (2009), as well as Escobar (2011), also show that the SDM subsumes not only the Spatial Error Model (SEM) and the Spatial Autoregressive (SAR) model, but also the spatial lag of X model (SLX) and OLS.

The general form of the SDM is:

\[ Y = \rho W Y + X \beta + WX \theta + \iota_n \alpha + \epsilon, \]

where \( Y \) is the dependent variable (quantity sold, in our study); \( W Y \) is the spatial lag of the dependent variable; \( \rho \) is the spatial lag coefficient for the dependent variable; \( X \) represents a vector of the explanatory variables of the model; \( \beta \) represents a vector of coefficient estimates, indicating the strength and direction of the impact of the explanatory variables on \( Y \); \( WX \) is the spatial lag of the explanatory variables; \( \theta \) is the spatial lag coefficient for the explanatory variables; \( \iota_n \alpha \) is a vector of constant terms; \( \epsilon \) is the error term, \( \epsilon \sim N(0, \sigma^2 I_n) \); and \( n \) is the number of observations.

Based on (7), the spatial Durbin model can be rewritten as:

\[ (I_n - \rho W) Y = X \beta + WX \theta + \iota_n \alpha + \epsilon. \]
Solving for $Y$ we obtain the data generating process for the SDM:

$$Y = \sum_{k=1}^{l} S_k(W) x_k + V(W) \xi_i + V(W) \varepsilon,$$

where

$$S_k(W) = V(W)(I_n \beta_k + W \theta_k)$$

and

$$V(W) = (I_n - \rho W)^{-1} = I_n + \rho W + \rho^2 W^2 + \rho^3 W^3 + \cdots$$

The powers of the matrix $W$ in $V(W)$ will form linear combinations based on neighbors of neighbors for each observation. Since the diagonal elements of the matrix $W^k$, $k > 1$, are not zero (units can be neighbors to their neighbors), a change in observation $i$ will determine an impact on second-order and higher-order neighbouring observations and back to $i$, or, in other words, feedback effects (LeSage and Pace, 2009, 2012). Expanding the relationship in (9) we obtain:

$$\begin{pmatrix} y_1 \\ y_2 \\ \vdots \\ y_n \end{pmatrix} = \sum_{k=1}^{l} \begin{pmatrix} S_1(W)_{11} & S_1(W)_{12} & \cdots & S_1(W)_{1n} \\ S_2(W)_{21} & S_2(W)_{22} & \cdots & S_2(W)_{2n} \\ \vdots & \vdots & \ddots & \vdots \\ S_n(W)_{n1} & S_n(W)_{n2} & \cdots & S_n(W)_{nn} \end{pmatrix} \begin{pmatrix} x_1 \\ x_2 \\ \vdots \\ x_n \end{pmatrix} + V(W) \xi_i + V(W) \varepsilon.$$  

(12)

Unlike linear regression, in the case of the SDM the derivative of $y_i$ with respect to $x_{ik}$ does not equal $\beta_k$ because of the presence of feedback effects and the derivative of $y_i$ with respect to $x_{jk}$, $\delta y_i / \delta x_{jk} = S_k(W)_{ij}$, takes a value determined by the $i,j$th element of the matrix $S_k(W)$ and it is potentially non-zero. This implies that a change in the explanatory variable (for example the price of a product) for a single observation can potentially affect the dependent variable (quantity sold) in neighbouring regions (competing ATC-codes), which represents the fundamental idea of spatial econometrics. Thus, as LeSage and Pace (2009) explain, the direct impact represents the impact from changes in the $i$th observation of $x_k$ ($x_{ik}$) on $y_i$ and it includes the feedback effects (general equilibrium effects in economic terms) where observation $i$ affects observation $j$ and observation $j$ also affects observation $i$. The indirect impact represents the impact on the $j$th observation's dependent variable $y_j$ from a change in the $i$th observation of the explanatory variable $x_k$ ($x_{ik}$). The total impact to an observation represents the impact on the individual observation $y_i$ from changes in the $k$th explanatory variable by the same amount across all observations (if averaging on columns). The total impact can also be read as the impact over all $y_i$ from changing the $k$th explanatory variable by an amount in the $j$th observation (if averaging on rows). The total impact less the direct impact yields the indirect impact. All the impacts are defined as averages in the sense that they represent means of the main diagonal (the direct effect) or of the off-diagonal (indirect effect) elements of the matrix $S_k(W)$ (LeSage and Pace, 2009). From the economic theory viewpoint, the direct effect is expected to be positive in our study as this is closely related to the own-price elasticity of demand, while the indirect effect is expected to be negative since this is similar to the cross-price elasticity of demand. Following Kirby and LeSage (2009), we can explain that the difference in interpretation of SDM coefficients as compared to an ordinary regression model such as OLS is that the latter would make the prediction that a change in the price of a product $i$, will affect only the quantity sold of product $i$, not allowing for spatial spillover impacts, while our model does.
TABLE 1
Variable definition

| Name      | Definition                                                                 |
|-----------|-----------------------------------------------------------------------------|
| **Dependent variable** | Natural logarithm of quantity sold (in units).                       |
| units     | Natural logarithm of price paid to the producer (per unit).                |
| **Explanatory variables** | Dummy variable equal to 1 if the product is parallel-imported from another European Union country, 0 otherwise. |
| price     | Natural logarithm of price paid to the producer (per unit).                |
| import    | Dummy variable defining product type:                                       |
|           | bgn 1 if product type is branded generic, 0 otherwise;                      |
|           | org 1 if product type is original (branded), 0 otherwise;                   |
|           | oth 1 if product type is different than the above and than generic, 0 otherwise. |
TABLE 2

Descriptive statistics of the dependent and explanatory variables

| Name   | Mean  | Std.Dev. | Min  | Max  |
|--------|-------|----------|------|------|
| units  |       |          |      |      |
| overall| 2.4636| 2.0210   | 0    | 10.9540 |
| between| 1.1841| 0.9439   | 10.0715 |
| within | 1.7324| 0.4658   | -1.4429 | 8.2698 |
| price  |       |          |      |      |
| overall| 4.4048| 1.3856   | -1.6094 | 10.0715 |
| between| 1.3786| 0.9439   | 10.0715 |
| within | 0.4658| -1.4429  | 8.2698  |
| import |       |          |      |      |
| overall| 0.0820| 0.2744   | 0    | 1    |
| between| 0.3273| 0.0064   | 0    | 1    |
| within | 0.0064| -0.0031  | 0.9987 |
| bgn    |       |          |      |      |
| overall| 0.1003| 0.3004   | 0    | 1    |
| between| 0.3145| 0.0394   | 0    | 1    |
| within | 0.0394| -0.8694  | 1.0675 |
| org    |       |          |      |      |
| overall| 0.3273| 0.4692   | 0    | 1    |
| between| 0.4540| 0.0193   | -0.6363 | 1.1145 |
| within | 0.0193| -0.6363  | 1.1145 |
| oth    |       |          |      |      |
| overall| 0.0209| 0.1430   | 0    | 1    |
| between| 0.1239| 0.0209   | 0    | 1    |
| within | 0.0209| 0.0209   | 0.0209 |

\[ \text{diff} = p(t) - p(t - 1), \text{where } p \text{ is the price per unit and } t, t - 1 \text{ are two successive time periods (not necessarily consecutive, since some products are not sold in all time periods).} \]

Figure 1 illustrates the statistically significant values of the Moran’s Is for price differences; those values that were not statistically significant show as zero in the graphs. According to the results, price changes for pharmaceuticals show a certain degree of positive autocorrelation. This autocorrelation is stronger at the seven-digit ATC level than at the other levels, meaning that the interdependence of changes in prices between close substitutes is, as expected, stronger. There is almost no correlation between prices of pharmaceuticals at the one-digit ATC level.

The values of the Moran’s Is are generally quite low, but mostly positive, except for a few isolate instances. There is a clear jump in these values in 2010, which could indicate a simultaneous decrease in prices as a response to the regulation changes implemented in 2009 and 2010. More insight into this matter is given by the analysis of the mutual interrelations between individual prices, which follows in the next section.

IV.3 Moran’s I associated with price per unit

As discussed, the Moran’s I statistics associated with price per unit help us determine the extent of the market for pharmaceuticals in Sweden. In a perfectly competitive market, the prices would be perfectly correlated with each other and thus the Moran’s I would be exactly 1 (e.g., Park and Kim, 2014).

We calculate the Moran’s I values for each product and time period. Figure 2 illustrates the results, which show significant dependence between prices of pharmaceutical products at all

\(^3\) The low Moran’s I values may be due to the fact that changes in prices as a result of price competition may occur in different time periods for different products (e.g., in subsequent time periods). By using the cross-sectional Moran’s Is we only investigate simultaneous price changes.

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ATC levels. However, for the three-digit and one-digit ATC levels the Moran’s I are lower than 0.3 and thus indicate that the law of one price clearly does not hold, and we interpret this as weak price competition. We can thus conclude that price competition on the pharmaceuticals market in Sweden extends to the four-digit ATC level inclusive. Figure 2 also indicates a reaction of the market to the policy changes that were implemented in 2009 and 2010 in the form of a turbulence that occurs in the beginning of 2010, when there is a sudden drop in the correlation of prices in the Swedish pharmaceuticals market at the seven-, five-, and four-digit ATC levels, which lasts for almost half year before the market recovers and the correlation suddenly increases to return to levels similar to the period before the implementations of the new regulations.

IV.4 Moran’s I and spatial lag coefficients associated with quantity sold

For the three ATC-code levels for which we have determined that the products belong to the same market, we further calculate the Moran’s I values associated with quantity sold. As the quantity sold at a specific period in time is mostly related to a product being the ‘product of the month’ and thus catering to approximately 80 percent of a given seven-digit ATC-code
exchange group (Tillväxtanalys, 2012), we expect that the Moran’s I for quantity sold will be low, and this is also what we find. However, although low (Figure 3), these Moran’s I values are still statistically significant and consequently they justify the use of a spatial model to determine the impact of prices on quantity sold.
This result is supported by the values of the scalar parameter $\rho$, the spatial lag coefficient for the dependent variable, which is another measure of the overall strength of spatial dependence between the observations in the dataset (Kirby and LeSage, 2009) and which can be estimated with the aid of any spatial autoregressive model (Figure 4). The bounds of the spatial lag coefficient are $-1/w_{\text{min}}$ and 1, where $w_{\text{min}}$ is the smallest characteristic root of the spatial weights matrix (Elhorst, 2010). Its interpretation is similar with that of the Moran’s I statistic: a positive spatial lag coefficient illustrates positive product space spillovers between the units of analysis as regards the variable of interest (quantity sold), while a negative rho parameter illustrates negative spatial spillovers between the units of analysis as regards the variable of interest. In our case, a positive and significant rho parameter associated with the dependent variable indicates that, to a certain extent, products registering high quantities sold are competing with other products registering high quantities sold, while products with low quantities sold are competing with other products with low quantities sold. However, the values of rho are quite low (generally lower than 0.40). They are also significant for most of the cross-sections, indicating the presence of spatial spillovers and supporting the use of a spatial regression model. A break occurs mid-2009, when the new regulations have been implemented, which determined the spatial lag coefficient for all three ATC levels analysed here to change from constant to slightly upward trending.

Following Brown et al. (2009), when rho is very small or close to zero the spatial Durbin model collapses to the spatial cross-regressive model (also known as SLX, the spatial lag of X)

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4The Moran’s Is do not take into account the effects of the independent variables as the estimate of $\rho$ does, and thus these are not directly comparable. However, as both the Moran’s Is and the results from the more formal statistical analysis show similar results, we consider this worth mentioning.
model), which can be estimated using OLS. However, if rho is statistically significant, results from the cross-regressive model will be biased and inconsistent (Anselin, 1988). In this case, it is still the SDM that provides more reliable estimates.

It must be mentioned here that the spatial lag coefficient, as well as the Moran’s I, are static measures, in the sense that they reflect steady state equilibriums. This means that if we have a high and significant Moran’s I value (or rho parameter, for that matter), we can argue that, at a certain point in time and to a certain extent (determined by the magnitude of this value), competing products have similar prices and sell in similar quantities. As expected, the relationship between prices of competing products is much stronger than the relationship between quantities sold of competing products due to the ‘product of the month’ system. However, based on these values we cannot tell how a change in the price/quantity sold of a product will impact the price/quantity sold of competing products or how the relationship between prices/quantities sold of competing products changes over time. An idea of the simultaneous variation in price changes is given in section V below where some causal interpretations are provided by the regression model in the form of the direct, indirect, and total effects.

V. RESULTS AND DISCUSSION

To determine the impact of price on quantity sold of pharmaceutical products in the presence of product space dependence, before and after the policy changes that took place in 2009 and 2010, we calculate the direct, indirect, and total price effects for the four-, five-, and seven-digit ATC levels (which according to our analysis define the extent of the market for pharmaceuticals in Sweden) and all available time periods (January 2006 to June 2011) on a bi-weekly basis. We use a Bayesian Markov Chain Monte Carlo method (MCMC) that provided 5,000 draws to calculate standard deviations from the posterior distribution of the effect estimates. The standard deviations are further used to calculate t-statistics and marginal probabilities on which inferences about the significance of the direct, indirect, and total effects are based. It has been previously shown that these draws can provide valid inferences in non-linear functions of these parameters (Gelfand and Smith, 1990; Gelfand et al., 1990; Kirby and LeSage, 2009).

The variables have been used in their demeaned form in the MCMC model to account for any time-invariant heterogeneity affecting both prices and demand for different pharmaceuticals. Given the highly-unbalanced panel we are dealing with, repeated cross-sectional analysis has proved a more appropriate form of analysis. However, by demeaning, the average values for each variable (both dependent and explanatory) are subtracted from the observed values of the respective variable. Consequently, within each time period, the mean of each demeaned variable is zero. This operation eliminates all between-periods variability and leaves in only the within-period variability to analyse (Studenmund, 2005). This is an innovation and a step forward from pure cross-sectional analysis towards panel data analysis. Results of the spatial Durbin model for one of the cross-sections (the second half of May 2011) are presented for exemplification in Table 3.

Reading through the direct impacts, we note that only those for price, import, and original product type are significant at the 95 percent confidence level at the seven-digit ATC code. At this level, a 10 percent increase in the price of a product determines a 3.11 percent decrease in the quantity sold of the respective product; parallel imported substitutes are sold at a 73.66 percent discount, while branded products imply a 36.91 percent higher price on average. On the other hand, increasing the price of the product by 10 percent determines a 1.62 percent increase in the quantity sold of competing products (close substitutes, in this case).

The results for the five- and four-digit ATC levels are similar. We note that the direct impact of a change in price at these levels is lower than at the seven-digit ATC level (a 10 percent
### TABLE 3
Results of the spatial Durbin model for the second half of May 2011 at the seven-, five-, and four-digit ATC levels

| Direct impact | Coefficient | Lower 05 | Upper 05 | Coefficient | Lower 05 | Upper 05 | Coefficient | Lower 05 | Upper 05 |
|---------------|-------------|----------|----------|-------------|----------|----------|-------------|----------|----------|
| price         | −0.3106**   | −0.3865  | −0.2356  | −0.2751**   | −0.3507  | −0.2012  | −0.2610**   | −0.3332  | −0.1920  |
| import        | −0.7366**   | −1.1969  | −0.2716  | −0.4968**   | −0.9396  | −0.0537  | −0.5744**   | −0.9688  | −0.1746  |
| bgn           | −0.0712     | −0.3305  | 0.1818   | −0.0686     | −0.3247  | 0.1781   | −0.0053     | −0.2382  | 0.2332   |
| org           | 0.3691**    | 0.1624   | 0.5742   | 0.4202**    | 0.2164   | 0.6214   | 0.4188**    | 0.2175   | 0.6147   |
| oth           | −1.5530     | −5.5935  | 2.5910   | −0.8223     | −1.8314  | 0.2340   | −0.8621     | −1.7799  | 0.0650   |
| Indirect impact | Coefficient | Lower 05 | Upper 05 | Coefficient | Lower 05 | Upper 05 | Coefficient | Lower 05 | Upper 05 |
| price         | 0.1617**    | 0.0281   | 0.2912   | 0.1073      | −0.0359  | 0.2506   | 0.0828      | −0.0844  | 0.2459   |
| import        | −0.6944     | −1.7632  | 0.3396   | −1.4587**   | −2.6711  | −0.2718  | −1.6038     | −3.3270  | 0.0360   |
| bgn           | 0.4829      | −0.0246  | 0.9973   | 0.5598      | −0.0208  | 1.1294   | 0.2115      | −0.4673  | 0.8663   |
| org           | 0.9405**    | 0.3867   | 1.5156   | 1.1249**    | 0.4354   | 1.8413   | 1.4531**    | 0.5395   | 2.4483   |
| oth           | 2.1429      | −2.1111  | 6.2141   | 1.8112**    | 0.3790   | 3.2207   | 2.0137**    | 0.6676   | 3.3550   |
| Total impact  | Coefficient | Lower 05 | Upper 05 | Coefficient | Lower 05 | Upper 05 | Coefficient | Lower 05 | Upper 05 |
| price         | −0.1489**   | −0.2703  | −0.0351  | −0.1679**   | −0.3047  | −0.0424  | −0.1782**   | −0.3359  | −0.0248  |
| import        | −1.4310**   | −2.4872  | −0.4617  | −1.9555**   | −3.1951  | −0.7688  | −2.1782**   | −3.9102  | −0.5285  |
| bgn           | 0.4116      | −0.0703  | 0.9015   | 0.4912      | −0.0604  | 1.0481   | 0.2062      | −0.4547  | 0.8568   |
| org           | 1.3096**    | 0.7443   | 1.8970   | 1.5451**    | 0.8564   | 2.2617   | 1.8719**    | 0.9495   | 2.8683   |
| oth           | 0.5904      | −0.1876  | 1.4279   | 0.9889**    | 0.0434   | 1.9828   | 1.1516**    | 0.0920   | 2.2517   |

**Significant at the 95 percent confidence level.
increase in price causes a 2.75 and 2.61 percent decrease in demand, respectively). The indirect effects of price are not significant any longer, which indicates that price competition is only present amongst close substitutes.

Figure 5 illustrates these direct, indirect, and total price effects on quantity sold for pharmaceutical products defined as competing on the basis of the seven-, five-, and four-digit ATC codes for all analysed cross-sections. The direct effect represents the marginal effect (the effect of a change in price on quantity sold of the same product) plus the feedback effect (the effect of a change in the price of a product on the quantity sold of the competing products, which in turn affects the quantity sold of the respective product). As expected, the direct effect,
which is closely related to the own-price elasticity of demand, is negative and significant for almost all the cross-sections included in analysis.\footnote{The difference is that usually the calculated own-price elasticities of demand do not contain feedback effects while the direct effects in our calculations do.}

For the seven-digit ATC code, the average of the direct effect over all time periods is $-0.22$ for those periods when the direct effect is significant. Figure 5 shows a slight drop in the direct effect after 1 July 2009, when the first new regulations were enforced, which indicates that an increase in price leads to a more considerable decrease in quantity sold after the reference point (1 July 2009) than before it. On average over all analysed time periods, the direct effect before 1 July 2009 is $-0.20$, while the average direct effect after 1 July 2009 is $-0.26$, which represents a 30 percent decrease, significant at the 99 percent confidence level. The feedback effect for the time period used for exemplification (the second half of May 2011) is $-0.0038$, which represents 1.73 percent of the direct effect. The feedback effects are calculated as the difference between the SDM coefficient estimate and the direct effect. Sometimes they are very small and not likely of economic significance (LeSage and Pace, 2009). In our study, feedback effects are quite low over the entire study period. The absolute value of the feedback effects over all analysed time periods is 0.0012 for those periods when the direct effect is significant (or 0.55 percent of the average direct effect).

The direct effect measures localized effects. The indirect effect measures the spatial spillovers. If both direct and indirect effects are significant we can argue that a change in price not only will impact the quantity sold of the own product, but will also spillover to the competitors. Following Escobar’s line of thought (2011) we can show that the induced changes in the competitor will further spillover to the competitors of the competitors, including the product in which the change originated, i.e., there are so called general equilibrium effects. This process of adjustments continues until a new steady-state equilibrium is reached (Molho, 1995; LeSage and Pace, 2009).

The indirect effect, or the product space spillovers, represents the effect of a change in own price on the quantity sold of the competing products - or the cross-price elasticity of demand. For the same time period considered above, the indirect effect is 0.16 measured at the seven-digit ATC code. This means that a 10 percent increase in the price of a product determines a 1.6 percent increase in the quantity sold of the competing products, on average for all products in this cross-section. It is interesting to note a slight upward trend of the indirect effect, especially after 1 July 2009, which shows a strengthening of competition at this ATC level: a 10 percent increase in the price of a product determines a higher increase in the quantity sold of competing products after 1 July 2009 than before this reference point. The average for the significant effects over all analysed time periods before 1 July 2009 is 0.04, while the average after 1 July 2009 is 0.16, which represents a 300 percent increase, significant at the 99 percent confidence level. As we will see in the following section, the indirect effect also becomes significant in many more of the analysed cross-sections after this date than before the new regulations have been introduced. Although the results for the seven-, five-, and four-digit ATC codes look similar in Figure 5, Figure 6 shows that there is an important difference between them in the significance levels of the indirect effect.

A look at the significance of the indirect effect offers a better idea of the changes the market for pharmaceuticals has gone through after the new regulations have been implemented. T-statistics have been calculated from a set of 5,000 simulated parameter values. Although not so visible at the 95 percent confidence level, when investigating the significance of the indirect effect at the 90 percent confidence level a clear shift is revealed from indirect effects being mostly insignificant before 2009 to indirect effects being mostly significant after 2009. The shift is observable at the five-digit ATC level and particularly at the seven-digit ATC level.
This result shows that the latest regulations have impacted these two levels in particular, in the sense that the market for close substitutes has started to work better and thus price competition has been strengthened. For the seven-digit ATC level, the indirect effect is significant at the 90 percent level for six cross-sections before 1 July 2009 and for 21 cross-sections after 1 July 2009 (a 250 percent increase). For the five-digit ATC level, the proportion is 6/9 (a 50 percent increase), while for the four-digit ATC level, the proportion is 8/6 (a 25 percent decrease). This is an indication that the policy changes from 1 July 2009 increased market effectiveness especially for those products that are close substitutes (the seven-digit ATC level).

The total effect represents the sum of the direct and the indirect effects and indicates the impact on an individual observation from changes in a particular explanatory variable by the same amount in all observations. The total effect is \(-0.15\) for the time period used for exemplification, indicating that a 10 percent increase in the price of all pharmaceutical products would cause a 1.5 percent drop in quantity sold on average for all pharmaceutical products sold during this time period. The total effect, averaged over all time periods before 1 July 2009 for which the effect was significant, is \(-0.19\), compared to the total effect, averaged over all time periods after 1 July 2009 for which the effect was significant, which is \(-0.16\). This indicates a 16 percent increase, significant at the 99 percent confidence level.

VI. CONCLUDING REMARKS

The contribution of this study to the literature on markets and competition is that it employs spatial regression methodology to determine, on the one hand, the extent of the product market for pharmaceuticals in Sweden based on price competition between pharmaceutical products, and on the other hand, how the policy changes implemented in 2009 and 2010 have affected the Swedish market for pharmaceuticals in terms of the impact of price changes on quantities sold of own and competitor products. To our knowledge, no previous study has addressed the issue of what is the relevant market for pharmaceuticals in Sweden. Given the particularities of the units of analysis (pharmaceutical products), our study also applies spatial econometric methodology on non-spatial items, which have been seldom performed previously.

Our results show that the relevant product market for pharmaceuticals clearly extends to the seven-digit ATC-code level where brand name products and generic copies with the same chemical substances compete with each other. There is also competition between different
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chemical substances belonging to the same chemical subgroup (five-digit ATC level), and we even find some weak evidence of competition on the pharmacological level (four-digit ATC level). As such, previous studies focusing on competition between brand name products and generics only might have had too narrow a focus, and thus underestimated the level of competition in the pharmaceuticals market.

Our results also indicate that the Swedish pharmaceuticals market is better functioning after the 2009–2010 reforms. An increase in the price of a product determines a higher increase in the quantity sold of competing products after 1 July 2009 than before the reform. The average of the indirect effects, representing the cross-price elasticity of demand, was 0.04 before 1 July 2009, while the average after 1 July 2009 is 0.16. This represents a 300 percent increase, significant at the 99 percent confidence level. Also, investigating the significance of the indirect effect before and after the reforms, a clear shift is revealed from indirect effects being mostly insignificant before the reforms to indirect effects being mostly significant after the reforms. The reform effect is observable at the five-digit ATC level and particularly at the seven-digit ATC level. This shows that the market reforms in 2009 and 2010 have impacted these two levels of the pharmaceutical market in the sense that the market for close substitutes has started to work better and price competition has been strengthened.

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