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### Abstract

Foreign manufacturers have the option of using sales networks of domestic rival firms to save local distribution costs. Such alliances may lead to collusion or create greater distortions because of the additional margins imposed by foreign firms, as shown in the theoretical literature. This paper empirically examines whether these outcomes are realized by alliances using Japanese antibiotics market data, where cross-border alliances are common. Empirical results show that the marginal costs of products supplied through cross-border alliances are lower than those supplied by foreign firms, suggesting that alliances are effective devices to reduce local distribution costs for foreign firms. Furthermore, my test results reveal little evidence of collusion or high markups caused by cross-border alliances.

Key words: Strategic Alliance; Market Competition, Pharmaceuticals; Discrete Choice Model

JEL classification: F23; L24; L65

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# 1 Introduction

There are effectively no tariff barriers for pharmaceutical products in Japan because of agreements related to the WTO. However, differences in laws, regulations, and distribution systems in health care and pharmaceutical markets may entail costs for foreign firms. In particular, the Japanese Pharmaceutical Affairs Act and local pharmaceutical distributions were considered to impose complex and costly processes on foreign companies. Thus, it is possible that foreign firms may form alliances with Japanese firms to alleviate such burdens to supply the Japanese market. In fact, drugs supplied through cross-border alliances are common in Japanese drug markets.

In addition to cost-saving motives, when firms compete in a product market, cross-border alliances can be used as collusive devices because firms are able to share profits through alliances. Thus, cross-border alliances can be considered anticompetitive; for example, distribution alliances in a highly concentrated market (e.g., Miller and Molson on Canada and the US beer markets in 1995) and code-sharing arrangements between airlines. Without collusion, because foreign firms may impose per-unit royalties on alliance products, higher margins are created because of double marginalizations, which then leads to market distortions (Tirole, 1988). While the possibilities of these anticompetitive effects of alliances have been studied theoretically (for international alliances, see Chen (2003), and for domestic alliances, see Katz and Shapiro (1985) and Fauli-Oller and Sandonis (2002)), whether the effects of cross-border alliances on product market competition exist is an empirical issue.

This study first addresses the issue of whether the supply costs of products through cross-border alliances are lower than those of foreign supply products. By using estimates of the demand for antibiotic drugs in Japan, I derive implied marginal costs, and compare the costs of alliance products with those of a foreign firm's own products. Then, I examine the effects of per-unit royalties on market distortions. Finally, I investigate whether the market is competitive or colluded. I empirically test whether collusion

schemes are more consistent with the data than competition schemes by using a similar method to that of Bresnahan (1987), Villas-Boas (2007), and Bonnet and Dubois (2008). My focus is close to that of Goldberg and Verboven (2001), in which study the comparison between collusion and competition was exercised by using pricing equations. This study contributes to the literature by revealing the supply condition differences between domestic and foreign firms in a market without trade barriers, and assessing whether market competition prevails when alliances are prominent. I use the discrete choice framework for demand estimation of Berry (1994), and consider pricing behavior in a product market. By using demand parameters, I derive the implied markups and marginal costs to evaluate the effects of alliances on market competition. While this type of framework is used to examine the impacts of trade policies in the trade literature (Goldberg (1995), Berry, Levinsohn, and Pakes (1999), Ohashi (2002), and Irwin and Pavcnik (2004)), to my knowledge there is no study of cross-border alliances. The determinants of cross-border alliances have been examined empirically (Fosfuri (2004)). However, such empirical studies have not explicitly taken market competition into account. Because theoretical studies focus on the relationship between alliance and market competition (for example, Katz and Shapiro (1985), Qiu (2006), and Ishikawa, Morita, and Mukunoki (2008)), this paper fills the gap between theoretical and empirical studies.

My estimation results show that the marginal costs of products supplied through cross-border alliances are lower than those of a foreign firm's own products. This suggests advantages in distribution for local manufacturers: if local firms have low distribution costs, originator firms supply through alliances, raise total profits, and then extract profit by imposing licensing fees. It may be costly for foreign firms to create effective local distribution networks so that their own supply costs are high, and those of local firms through alliances are low. Thus, this study contributes to the literature by revealing the empirical determinant of cross-border alliances: the cost saving motive.

During my sample period, the marginal cost differences between domestic supply products, foreign supply products, and alliance products were likely to decrease, and in particular foreign supply costs decreased. This corresponds to the period when the Pharmaceutical Affair Act was revised, foreign subsidiaries were reorganized, and foreign firms' distribution strategies changed. The main aim of the Pharmaceutical Affair Act revision was to deregulate the approval system for drug sales. Before the revision, all firms selling drugs presumably had manufacturing plants, so the government imposed manufacturing approval for sales. This revision meant that firms selling drugs had to be authorized by the government. Because the approval system was similar to those in the EU and the US, foreign manufacturing firms' disadvantages may have been reduced.

Reorganizations of foreign subsidiaries, such as mergers, may also create effective local distribution channels in Japan. It has been argued that supplying drugs effectively in the Japanese market requires sufficient numbers of medical representatives (MRs). An MR is a sales person who promotes drugs to doctors. Promotion to doctors by MRs is considered an essential sales channel in Japan. By reorganizing subsidiaries, foreign firms can employ sufficient MRs. For example, Abbott's new subsidiary was created by merging Dinabott, with 400 MRs, and Hokuriku Seiyaku, with 300 MRs, which amounts to approximately 30 percent of the third-largest Japanese pharmaceutical company, Astellas. Thus, scale merits in distribution activities may generate cost reductions. In fact, in the later years in my sample, several alliance contracts were terminated and foreign originator firms began to supply independently.

By using demand estimates, I examine the relationship between royalty rates and pricing behavior. The royalty rates are not observed in the data, therefore I conduct hypothetical examinations by setting several royalty rates and deriving equilibrium behavior. My numerical examinations suggest that with royalty payments, high royalty rates are consistent with high markups. Thus, if high royalty rates are imposed, the anticompetitive effect of alliances may be severe. In addition, collusion behavior is

consistent with profit maximizing behavior, because implied marginal costs turn out not to be negative. As theoretical studies such as that of Chen (2003) demonstrate, the possibility of collusion caused by cross-border alliances may have adverse effects on market competition.

However, by examining the fit of functions of marginal costs, empirical tests show that collusion schemes are rejected in the case of Bertrand competition with only fixed-fee alliances. High per-unit royalty rate schemes are rejected by collusion schemes. The scheme that is never rejected is an alliance with fixed fees. This suggests that royalty rates may be zero and fixed fees are used, and alliances are not used as a collusive device. Hence, the anticompetitive effects of alliances on competition caused by collusion or high markups are not found. While, as theory suggests, the pricing behavior can be manipulated by royalties (Shapiro (1985)) and alliances can be used to mitigate intensive market competition (Gallini (1984), Rockett (1990), Eswaran (1994), and Fauli-Oller and Sandonis (2002)), market competition prevails in this market. This paper provides empirical findings on the effect of cross-border alliances and insight into the evaluation of competition in a market when alliances are active.

The paper is organized as follows. Section 2 introduces the data set. In Section 3, I set up the model and in Section 4, estimation results are reported. In Section 5, several alliance and collusion schemes are examined, and the final section concludes the paper.

## **2 Data and Background**

In this section, I introduce my sample data. I specifically consider the antibiotic drug market. This category includes injectable and oral antibiotics: the antibiotics beta-lactam, cephalosporin, glycopeptide, macrolide, and the quinolone antibacterial drugs. The data on approximately 30 drugs for the period 1999–2006 yield an unbalanced panel sample of 241 antibiotics. There are 17 firms in my sample, of which seven are foreign. For example, GlaxoSmithKline (GSK), and Abbott sell drugs on their own

and also form alliances with Japanese companies. On the other hand, Bristol–Myers Squibb (BMS) and Pfizer only sell drugs independently. In my sample, because five of seven alliance products are supplied through cross-border alliances, the effects of cross-border alliances will be prominent. One product, “Banan,” was introduced by both the originator, Daiichi Sankyo, and GSK. It is supplied by a foreign firm, but the originator is a Japanese company.

By focusing on pharmaceutical firms that compete horizontally in Japan, the competition effect of alliances is examined. For pharmaceutical products, it is easy to ascertain whether a particular drug is supplied by the originator or a partner firm (licensee). If a drug is supplied by the partner firm, we can learn not only the name of the distributing partner but also that of the originator firm. Therefore, I can identify any differences that exist in supply conditions and pricing behavior between a firm’s own and alliance products in the pharmaceutical market.

There are effectively no trade barriers in Japan because of agreements related to the WTO. However, the differences in regulations between Japan and other countries were recognized and considered to impose possible entry barriers. The Japan External Trade Organization (JETRO) published “The Survey on Actual Conditions Regarding Access To Japan – Pharmaceuticals” concerning this issue in 1998. This report was released immediately before my sample period; thus, the changes in regulations and market environment made during my sample period reflect its influence. The main point related to distribution activities is that approval systems vary among the EU, Japan, and the US.

As in the EU and the US, pharmaceutical products require approval for sale. However, the necessary approval license is different in Japan. In the EU and the US, firms selling drugs on the market require approval. In contrast, in Japan pharmaceutical firms supplying drugs presumably own manufacturing plants, and therefore the firms that sell drugs were expected to be approved as manufacturers. Thus, there was no possibility of outsourcing entire manufacturing processes and selling drugs. In addition,

importers “must obtain an import and sales business license from the Minister of Health and Welfare”<sup>1</sup> and had to obtain approval for each product. In addition, if importers sold drugs to medical institutions, they had to obtain a wholesale business license. In contrast, no license is required for import into EU countries and the US. Thus, before 2005, approval at multiple levels was required, and was thus considered to impose additional costs. Under the 2005 revision of the Pharmaceutical Affair Act, distributors need to be authorized, while manufacturers not engaging in distribution activities do not need approval. This approval system is similar to those of the EU and the US. Importer licenses are no longer required. Authorized distributors can import drugs on the Drug Master File (DMF) list. Thus, distribution costs are saved by eliminating multiple business licenses and regulations about pure manufacturing factories.

Another important point in the Japanese market is price regulation. Retail prices are regulated and set by the Japanese government. However, because wholesale prices are determined by pharmaceutical companies, there is still price competition in wholesale prices. Thus, I examine pricing behavior in wholesale markets. The vertical structure in Japanese pharmaceutical distribution is such that manufacturers must first sell drugs to wholesalers, which then supply them to hospitals. Before 1991, when the Japanese antimonopoly laws were amended, manufacturers could control the prices that wholesalers were able to set for hospitals. Such exclusive exercises are now prohibited. Then, while the manufacturers determined a single price for their wholesalers, the prices to hospitals were determined by the wholesalers. We consider the prices when manufacturers sell drugs to their wholesalers. Because a substantial proportion of doctors prescribe and dispense drugs, they may care not only about the wholesale price, but also about retail (dispensed)–wholesale price margins. Because this stimulates demand for drugs, as Iizuka (2007) showed, it may be necessary to take physician margins into account.

Iizuka (2007) derives the wholesale price using a regulatory formula for retail prices. The pricing

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<sup>1</sup>From “The survey on actual conditions regarding access to Japan pharmaceutical,” page 1, published by JETRO.

formula for retail price at year  $t$  is given by the regulatory form every two years:  $P_t = W_{t-1} + RP_{t-1}$ , where  $P_t$  is the retail price,  $W_{t-1}$  is the wholesale price collected and calculated by the government, and  $R = 0.02$ , set by the government. Hence, we can derive wholesale prices from  $W_{t-1} = P_t - RP_{t-1}$ . While, as in Iizuka (2007), the price data can be derived from this formula, I use the actual wholesale prices in a particular wholesale market in Tokyo, Kanda. The drug price set by manufacturers is called the bulkhead price. However, based on sales volumes and information on drug sales provided to manufacturers, rebates are paid to wholesalers. Thus, the actual price will be lower than the bulkhead price. The market in Kanda can be regarded as one between wholesalers. The transactions in this market are made in cash, and drugs are supplied and demanded at prices at which rebates may be taken into account. Therefore, the prices in this market may reflect actual trading wholesale prices. I use the price of a particular type of form (capsule, tablet, or granule) reported by the publisher of Information on International Pharmaceuticals. Two products, Banan and Modacine, are supplied by both the originator and the partner firm. Prior to 2003, the partner firms' prices were not reported. Because in other periods the prices of the originator and partner firms are often the same, I set the pre-2003 prices at the same level. In addition, the type of delivery of one drug, Pansporin, was changed so that the wholesale price is adjusted by the associated ratio of retail prices. In the estimations, the prices are deflated by the producer price index.

Because my sample includes many forms of drugs, prices need to have a common basis. As in the previous study (Stern (1996)), I use patient per-day price. Per-day prices for patients are derived by calculating the required dose multiplied by the unit price of each drug based on the recommended dosage. I then use sales of each drug to obtain quantity. Dividing sales by per-day price yields quantity, which is the number of patients in each year. Then, the number of patients is divided by the total number of patients to derive market share of each product. Because I have no data on the likelihood of diseases,

I use several measures of total number of patients. Total market size is from the Patient Survey by the Japanese government, and uses the estimated number of patients in Japan. Because I classify two groups of drugs (oral and injectable), the within-group shares are also calculated.

One feature of my data is that it does not capture all products in each market. However, our data cover the majority share in the market. For example, in 2004 the sample includes more than 90 percent of the market share of sales in this therapy area (the total value of sales of the sample drugs is 324.3 billion yen, while that of total sales in this therapy area is 350.6 billion yen). This may be sufficient for estimating strategic behavior by large companies. In addition, my data do not cover generic drugs, because they are still not widely available in Japan, and my focus is on the alliance activities in brand drugs. As long as there are benefits from alliances, generic entry may not drastically change supply mode choice. In fact, some drugs supplied through alliances (for example, “Kefral” is supplied by Shionogi and the originator is Eli Lilly) face generic drug entry, but other drugs supplied by independent firms also face generic entry (for example, “Ciproxan” supplied by Bayer). This fact also supports the view that firms do not simply form alliances for patent expired products, nor do they supply patented drugs independently.

== Table 1 Here ==

Table 1 reports summary statistics for drugs. The oldest drug was launched in 1971 and the newest in 2002. Half-life is an indication of a drug’s characteristics, and is the time required for a drug concentration to be reduced by 50 percent. Indication is the sum of the number of approved indications and the number of target bacteria. Contraindication is the number of contraindicated conditions. Foreign is a dummy variable, which takes the value of 1 when foreign companies supply the drug. These characteristics are used in Iizuka (2007). By comparing the oral and injectable drug prices, per-day price varies widely for injectable drugs.

The differences in the prices of drugs obtained through alliances and by a firm's own supply are shown in Table 2. The simple average price of a firm's own drug is lower than that for one obtained through alliances for oral drugs, but this is higher for injectable drugs. Thus, firms do not simply supply low-price drugs through alliances. However, their own and alliance products can differ in market penetration as a consequence. If their own supply has an advantage in obtaining large market share compared with supply through alliance, share-weighted prices will be higher for their own products than for the products of alliances. The lower half of Table 2 reports share-weighted prices. The weighted average prices of firms' own products are higher than those of alliance products. This implies that alliance products tend to have low market share, while that of firms' own products is higher. The market share, i.e., market penetration, differs among distribution channels, so the choice of distribution channel is critical for raising revenues in this market.

== Table 2 Here ==

This fact is consistent with foreign firms' strategies in recent years. Foreign firms reorganized their subsidiaries by merging with other firms and establishing distribution subsidiaries during my sample period. Once foreign firms can establish sufficient distribution ability, their sales strategy changes from alliance to direct supply. There was a trend for cross-border alliance contracts to be terminated, and foreign firms tended to supply the drugs independently.

The estimations use pooled data over a period of years. The factors of consumer choice specific to each year are controlled by year dummies. While the prescribing behavior may not change drastically over time, time-specific shocks can be taken into account by time-specific effects.

### 3 Model

I estimate a discrete choice model of demand developed by Berry (1994). Pharmaceutical demand estimation is considered within this framework by Stern (1996) and Iizuka (2007). While retail prices are regulated in the Japanese drug market, wholesale prices are set by pharmaceutical companies. Choice of drug is made by doctors, and my concern is their sensitivity to wholesale prices when prescribing drugs. In Japan, a large number of doctors prescribe and dispense drugs to their patients, and thus doctors can profit from retail–wholesale price margins. Although the separation of prescribing and dispensing has been pursued, even at the end of my sample period in 2006, about 45 percent of drugs were dispensed by doctors.<sup>2</sup> In addition, as Iizuka (2007) shows, while the Japanese doctors may care about the costs to patients when they prescribe drugs, their decisions are influenced by the margins between retail and wholesale prices. Thus, doctors may be sensitive not only to wholesale prices, but also to physician margins. This element is also taken into account when considering drug demand.

We assume that doctor  $i$ 's problem is to choose drug  $j$  among  $J + 1$  alternatives to maximize the indirect utility function:

$$\max_j V_{ij} = x_j\beta + \alpha p_j + \xi_j + \mu_{ij},$$

where  $x_j$  is product characteristics,  $p_j$  is price,  $\xi_j$  is unobservable product characteristics, and  $\mu_{ij}$  is an error term.  $x_j$  may include physician margins. Because retail and wholesale prices are highly correlated in the data, simply including retail price causes a multicollinearity problem in the estimations. Instead, markup rates, i.e., retail minus wholesale prices divided by wholesale prices, are introduced. Let  $\delta_j$  denote mean utility,  $\delta_j = x_j\beta + \alpha p_j + \xi_j$ . The mean utility for outside goods is normalized to zero,  $\delta_0 = 0$ . The error term has a group-specific component:  $\mu_{ij} = \eta_g(\sigma) + (1 - \sigma)\epsilon_{ij}$ .  $\sigma$  is a measure of the

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<sup>2</sup>Japan Pharmaceutical Association website, <http://www.nichiyaku.or.jp/index.html> (in Japanese).

group-specific effect, showing the within-group correlation. In my sample, there are two types of drugs in a broad sense, oral and injectable. Thus, I divide drugs into these two groups.

Assuming that the error term follows extreme distribution, the market share is expressed by a nested logit form (see Berry (1994)):

$$S_j = e^{\delta_j/(1-\sigma)} / D_g^\sigma (\sum_g D_g^{1-\sigma}),$$

where  $D_g = \sum_{j \in g} \exp(\delta_j/(1-\sigma))$ . We estimate the following equation:

$$\ln S_j - \ln S_0 = x_j \beta + \alpha p_j + \sigma S_{j|g} + \xi_j,$$

where  $S_{j|g}$  is within-group share. Because price, physician markup, and within-market share are endogenous, we require an instrumental variable estimation.

Following Berry, Levinsohn, and Pakes (1995) and Iizuka (2007), the sum of product characteristics for other firms' products and product characteristics of other products sold by the firm are used as instruments. Because of oligopolistic interaction among pharmaceutical firms, the presence of substitutable goods affects the wholesale price. In addition, because multi-product firms choose prices to maximize total profits, their prices depend on the characteristics of their other products. I also use information on foreign supply and competition intensiveness.<sup>3</sup> In the pharmaceutical industry, product characteristics are determined by research processes. This stage is before the pricing decision stage; therefore, product characteristics are considered to be exogenous.

My principal assumption is that alliance choices are made before pricing decisions. In other words, alliances are exogenous when firms set prices. This holds well if firms' decisions on forming alliances

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<sup>3</sup>The instruments are the sum of drug age, number of indications, number of contraindications, and half-life. Another characteristic of a product supplied by an independent firm is the ratio of other products' half-lives to their own product's half-life. The number of foreign firms and number of products are also used.

takes place before pricing decisions. How alliance agreements are determined is an interesting question, but is beyond the scope of this paper.

First, firms are assumed to compete in a static Bertrand fashion. Collusion cases are considered later. The firms maximize their profits by choosing the prices of each drug in each year. In general, firms produce multiple products. Thus, the profit functions for an originator firm  $k$  and a partner firm  $k'$  without per-unit royalties are, respectively:

$$\pi_k = \sum_j (p_j - c_j) S(p_j, p_{-j}) M + F,$$

$$\pi_{k'} = \sum_j (p_j - c_j) S(p_j, p_{-j}) M - F',$$

where  $F$  and  $F'$  are total fixed fees. For the originator firm, this  $F$  is not cost, but contributes to revenue. This is assumed to be sunk when firms determine their prices. Then, the first-order condition for product  $j$  is:

$$\sum_i (p_j - c_j) (\partial S_i / \partial p_j) + S_j = 0.$$

In matrix notation,  $\Omega(p - c) + S = 0$ , where  $p$  is the price vector,  $c$  is the marginal cost vector,  $S$  is the share vector, and  $\Omega$  is the matrix of price derivatives of market share. In the nested logit model, the derivatives of the market share with respect to price (see Berry (1994) and Ohashi (2002)) are expressed by:

$$\partial S_i / \partial p_j = \begin{cases} \alpha S_j (1 - \sigma S_{j|g} - (1 - \sigma) S_j) / (1 - \sigma) & \text{if } i = j \\ -\alpha S_j (\sigma S_{i|g} + (1 - \sigma) S_i) / (1 - \sigma) & \text{if } i \neq j, i, j \in g \\ -\alpha S_i S_j & \text{if } i \in g, j \in g', g \neq g'. \end{cases}$$

The price–cost margins are implied by the first-order conditions:  $p - c = \Omega^{-1} S$ . Similarly, marginal costs can also be derived:  $c = p - \Omega^{-1} S$ . In the pharmaceutical industry, R&D costs are the major costs, and are sunk when supplying drugs, so that marginal costs of production may be negligible. In fact,

R&D spending usually accounts for about 10 percent of revenues for large pharmaceutical companies. However, from financial reports of pharmaceutical companies, not only production but also distribution costs take a significant share of revenues, so my implied marginal costs reflect these.<sup>4</sup>

One remark should be made here. Although my focus is on the differences in local supply costs between domestic and foreign firms, that between distribution and production costs is not identified. It may happen that while distribution costs of domestic and foreign firms are similar, foreign production costs are higher, and thus foreign supply costs appear to be higher than domestic supply costs. Therefore, our study investigates overall supply costs, including distribution and production costs.

With respect to royalty payments, I simply assume that a portion of sales is paid to originator firms. Let  $\lambda$  denote the royalty rate, therefore net profits for a licensee are given by:

$$\pi_k^\lambda = \sum_j (p_j - c_j) S(p_j, p_{-j}) M - F' - \sum_m \lambda p_m S_m M, \quad (1)$$

where the subscript  $m$  indicates licensed products. Thus, partner firms discount their sales by  $\lambda$ . They do not take  $p_j S_j M$  as revenue from product  $j$ , but take  $(1 - \lambda) p_j S_j M$  as a base for revenues. At the same time, originator firms receive  $\lambda p_j S_j M$  as royalty payments. For simplicity, the same  $\lambda$  is applied to all alliances in a market.

One important issue here concerns the reality of the royalty payment form I use. In the literature, different licensing contract forms, such as fixed-fee only, royalty only, and two-part tariffs, are reported (Macho-Stadler, Martinez-Giralt, and Perez-Castrillo (1996)). While the details of alliance contracts are not generally publicly available, we have supplemental evidence about alliance contracts in which US pharmaceutical firms are involved. In the US, public companies must file a report to the SEC about important contracts. Hence, supporting examples of alliance contracts between big pharmaceu-

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<sup>4</sup>Chaudhuri, Goldberg, and Jia (2005) derive upper and lower marginal costs of Indian pharmaceutical firms considering perfect competition and monopoly, respectively.

tical companies may be found. Basically, both fixed-fee and per-unit royalty alliance contract types are found. For example, the report of the alliance contract between Genelabs and Tanabe (a Japanese company, now Mitsubishi–Tanabe) documents that the royalty payments are determined by the following per-unit royalty rule: “\*\* percent of the portion of annual Net Sales for such calendar year which are equal to or less than U.S. \*\* dollars” (\*\* are sealed).<sup>5</sup> This alliance covers broad activities such as patent licensing, know-how transfer and commercialization.

The use of both fixed fees and royalty payments is found. For example, the copromotion agreement between Santarus and Otsuka includes upfront payments, which correspond to our fixed fees, milestone payments, which are payments contingent on market sales, and royalties. The royalties are specified as follows: “an amount equal to [\*\*\*] percent [\*\*\*] (the “Royalty”) multiplied by Net Sales”, where “Net Sales” correspond to our sales variable.<sup>6</sup> Therefore, both fixed-fee and per-unit royalty rates are used. My specification of alliance contracts can be considered a simplified version of such payment schemes. Thus, my specification that there is a fixed fee,  $F$ , and  $\lambda$  portion of sales is paid to the licensor simplifies the actual alliance contracts among pharmaceutical companies, but is not unrealistic. While some agreements may be based on R&D spending, if these costs are proportional to sales, my specification does not deviate greatly.

To provide an idea of how equilibrium behavior changes according to form of payment, consider an example of three firms and four products: products 1 and 2 are supplied by firm 1, product 3 is supplied by firm 2, product 2 is licensed to firm 1 by firm 2, and product four is supplied by firm 3. Then, the

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<sup>5</sup>This document is available at <http://www.sec.gov/Archives/edgar/data/874443/000095013404006962/f98659exv10w17.txt>.

<sup>6</sup>This is available at <http://www.sec.gov/Archives/edgar/data/1172480/000095013704009765/a03105exv10w5.txt>.

first-order conditions under per-unit royalty rates are given by:

$$\begin{pmatrix} \partial s_1/\partial p_1 & \partial s_2/\partial p_1 & 0 & 0 \\ \partial s_1/\partial p_2 & \partial s_2/\partial p_2 & 0 & 0 \\ 0 & 0 & \partial s_3/\partial p_3 & 0 \\ 0 & 0 & 0 & \partial s_4/\partial p_4 \end{pmatrix} \begin{pmatrix} p_1 - c_1 \\ (1 - \lambda)p_2 - c_2 \\ p_3 - c_3 \\ p_4 - s_4 \end{pmatrix} + \begin{pmatrix} s_1 \\ (1 - \lambda)s_2 \\ s_3 + \lambda p_2 \partial s_2/\partial p_3 \\ s_4 \end{pmatrix} = 0.$$

In general, using matrix notation, this is expressed by:  $\Omega(Jp - c) + \tilde{S} = 0$ , where  $\tilde{S}$  is  $S$  with partial derivatives included if the product is licensed out, and  $J$  is the identity matrix whose  $(i, i)$  element is  $1 - \lambda$  if the  $i$ th product is obtained through an alliance. Therefore, the implied marginal costs are given by  $c = Jp - \Omega^{-1}\tilde{S}$ . Given different values of  $\lambda$ , the corresponding implied marginal costs are derived. These implied marginal costs yield the margins  $p - c$ . Note that because of common  $\lambda$  the effects of royalty rates should be interpreted as market average effect.

The market competition mode has been assumed to be a static Bertrand. There is no information on the actual competition mode. Therefore, I also consider collusion schemes, and derive firm behavior. Two types of collusion are considered. One is a fully collusive scheme, where a hypothetical monopolist maximizes the joint profits, and the other is partial collusion, in which firms forming alliances are considered single firms.

The full collusion case is extreme. All firms in this market collude. Because, in this market, more than 30 percent of products are supplied through alliances, if the prominence of alliances leads to collusive behavior in a whole market, this full collusion scheme may arise. The hypothetical monopolist's

profit maximization condition in the above example is expressed by:

$$\begin{pmatrix} \partial s_1/\partial p_1 & \partial s_2/\partial p_1 & \partial s_3/\partial p_1 & \partial s_4/\partial p_1 \\ \partial s_1/\partial p_2 & \partial s_2/\partial p_2 & \partial s_3/\partial p_2 & \partial s_4/\partial p_2 \\ \partial s_1/\partial p_3 & \partial s_2/\partial p_3 & \partial s_3/\partial p_3 & \partial s_4/\partial p_3 \\ \partial s_1/\partial p_4 & \partial s_2/\partial p_4 & \partial s_3/\partial p_4 & \partial s_4/\partial p_4 \end{pmatrix} \begin{pmatrix} p_1 - c_1 \\ p_2 - c_2 \\ p_3 - c_3 \\ p_4 - s_4 \end{pmatrix} + \begin{pmatrix} s_1 \\ s_2 \\ s_3 \\ s_4 \end{pmatrix} = 0.$$

The partial collusion case is more reasonable than the full collusion case. If, in fact, firms in alliance collude, their decisions may resemble those of a single firm. In my example, if firms forming alliances maximize their joint profits, firms 1 and 2 are considered a single firm, and products 1, 2, and 3 are considered to be supplied by that firm. Therefore, the FOC is:

$$\begin{pmatrix} \partial s_1/\partial p_1 & \partial s_2/\partial p_1 & \partial s_3/\partial p_1 & 0 \\ \partial s_1/\partial p_2 & \partial s_2/\partial p_2 & \partial s_3/\partial p_2 & 0 \\ \partial s_1/\partial p_3 & \partial s_2/\partial p_3 & \partial s_3/\partial p_3 & 0 \\ 0 & 0 & 0 & \partial s_4/\partial p_4 \end{pmatrix} \begin{pmatrix} p_1 - c_1 \\ p_2 - c_2 \\ p_3 - c_3 \\ p_4 - s_4 \end{pmatrix} + \begin{pmatrix} s_1 \\ s_2 \\ s_3 \\ s_4 \end{pmatrix} = 0.$$

I derive marginal costs under competition and collusion schemes and regress the marginal costs on product and firm characteristics. We cannot detect competition modes separately from marginal costs because of the data limitation of marginal costs. Rather, the estimates of functions under collusion and alliances are compared to determine the function that best fits the data, and then the best-fit case is chosen as the prevailing competition mode.

The comparison between competition and collusion using pricing equations has been conducted by Goldberg and Verboven (2001). In this paper, I use a similar approach to Bresnahan (1987), Villas-Boas (2007), and Bonnet and Dubois (2008) to test whether this market is competitive or colluded under alliances by fitting functions to marginal costs. Because the estimated parameters in functions may be biased without cost data, as Corts (1999) and Genesove and Mullin (1998) show, I do not provide an interpretation of each parameter, but examine the fit of each model. I use the Vuong (1989) test to

examine which competing model fits the data best. One useful aspect of this test is that each model can be misspecified, which may hold in our case.

## 4 Results

### 4.1 Demand Estimation

Table 3 reports demand estimation results. Columns 2 and 3 report logit demand estimation results, and columns 4 to 7 report nested logit demand estimation results.

In column 2, OLS estimation results show that the coefficient of price is negative and significant. Therefore, doctors are sensitive to drug prices, and thus a drug demand function is found. In column 3, the IV estimation results of the same specification are reported. The price coefficient is also significantly negative. A comparison of columns 2 and 3 shows that the price coefficient is underestimated in OLS. This finding is consistent with the case that prices are correlated with unobservable product characteristics.

Similar results are obtained in the nested logit estimations. Columns 4 and 5 report the results of nested logit estimations. The price coefficients are significantly negative. In the nested logit estimations, the price coefficient is also underestimated in OLS. The within-share coefficients are about 0.54 and significant. Therefore, the drug-type choice is consistent with consumer behavior, and significant for drug demand. The magnitude of the within-share coefficient is smaller in OLS than in IV. The similar bias to price coefficient estimations exists in within-share estimations. The overidentifying restrictions test is satisfied ( $\chi^2(3) = 5.021$  and the p-value is 0.17), therefore the instruments are valid.

Columns 6 and 7 in Table 3 report the results when physician margins are included. The estimation results are qualitatively similar to other results. While the coefficient of wholesale price is negative, that of markup rate is positive in the IV estimation in Column 7. This may indicate that physicians'

demand behavior depends on the wholesale price negatively and on their margins positively, however the effect of physician margins is statistically insignificant. The coefficients of product characteristics reflect consumer preferences for antibiotic drugs in Japan.<sup>7</sup>

Because this is a semi-log linear demand model, the marginal effect of a 1 percent price increase is  $0.07 \times 15.1 = 1.057$  when evaluated at the average drug price, 15.1 (in 100 yen). Because retail prices are regulated, wholesale price competition must be restricted by regulation. Therefore, the demand function is not very elastic with price.

== Table 3 Here ==

## 4.2 Implied Marginal Costs and Markups

From the demand estimations, I derive markups and marginal costs. My focus is on the average differences at the market level between products obtained through alliances and firms' own products. Therefore, I examine aggregate markups using a weighted average of markups (see for example, Ohashi (2002)). The market-level markups are calculated as follows:  $s'(p - c) = s'\Omega^{-1}S$ , where  $s$  is the within-drug supply pattern (alliance, firm's own, or foreign supply) share,  $\sum_{j \in f} s_j = 1$ ,  $j$  denotes each drug, and  $f$  denotes each supply pattern. The weighted average of markups in each year is calculated by using shares as weights.

To show the statistical significance of my results, Figures 1 and 2 display the fifth and 95th percentile points of marginal costs and markups, respectively, by using the Monte Carlo method. I generate 100 draws from a normal distribution for the demand estimates to construct a 90 percent confidence interval.

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<sup>7</sup>My demand estimation is similar to that of Iizuka (2007), in which the demand for hypertension drugs is estimated. While the coefficients of the indication and half-life variables show the opposite sign to those in Iizuka (2007), this may be because these studies are concerned with different classes of drugs. For antibiotics, because of resistant bacteria, the number of indications and half-lives may not reflect the quality of the drugs.

Figure 1 shows that marginal costs of products supplied through cross-border alliances are lower than those supplied by foreign firms. This implies that local manufacturing firms may have a cost advantage in distribution, and reveals that a firm's alliance strategies are based on efficiency. If partner firms have low costs, it is beneficial for originator firms to supply through alliances and raise profits by imposing licensing fees.

== Figure 1 Here ==

The cost differences between domestic supply, alliance, and foreign supplies are not caused by price differences alone. As Table 2 reports, the average price of domestic supply products is lower than that of foreign supply and alliances, except for foreign alliance injectable drugs. Because prices reflect drug quality to some extent, foreign firms do not simply outsource low-quality products and keep high-quality goods for themselves. In addition, the share-weighted prices of foreign supply and alliance products are lower than of those supplied domestically. Therefore, the distribution channel affects market share, and thus market penetration.

Figure 1 also shows that as years pass, marginal cost differences narrow, and in particular costs of foreign firms' own supply decrease<sup>8</sup>. This salient feature reflects the changes in regulations and foreign supply conditions. There was a revision of the Japanese Pharmaceutical Affair Act in 2005. As discussed above, distributors need to be authorized, and pure manufacturers do not require approval for sales. Thus, foreign firms face a similar regulatory system to that in their home countries. In addition, the business license system was simplified. There is no longer an importer license, and authorized distributors can import drugs on the DMF list. This deregulation may reduce foreign firms' supply

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<sup>8</sup>Note that because one in seven product alliances is between domestic firms, a similar computation using not only cross-border but also domestic alliance products applies. The results of marginal costs of domestic alliances and foreign alliances are shown not to be significantly different. Therefore, my results imply that domestic own supply is the most efficient method for sales.

costs.

The foreign firms' cost changes may also be because of their subsidiary restructuring and distribution strategies. Around 2002 and 2003, the subsidiaries of foreign pharmaceutical companies in Japan were reorganized. Because of the merger between Pfizer and Pharmacia, the Japanese subsidiaries were also merged in 2003. In the same year, Abbott Japan was established by merging Dainabot (Abbott and Dainippon Seiyaku JV) and Hokuriku Seiyaku. BMS built a distribution subsidiary in 2002. This type of restructuring occurred in the Sanofi Aventis and GSK subsidiaries in Japan (GSK's Japanese subsidiary became a 100 percent subsidiary in 2006 and Aventis Pharma KK and Sanofi Synthelabo KK merged to create Sanofi Aventis KK in 2006 in Japan). Merck had a 49 percent share of a Japanese pharmaceutical company, Banyu, and acquired full ownership in 2004. Therefore, foreign firms may reduce the disadvantage of supplying their own drugs in Japan.

Associated with foreign subsidiary reorganizations, foreign firms' distribution strategies have changed. It has been documented that foreign firms with distribution ability began to terminate alliance contracts and supply drugs by themselves to a greater extent than previously. In 2006, the originator firms of two alliance products, Lulid and Modacine, Sanofi Aventis and GSK, became the suppliers. Thus, foreign organizations and strategies were drastically changed at this time. As mentioned in the introduction, promotion of pharmaceuticals to doctors by MRs is an important sales channel in Japan. Reorganizing subsidiaries yields a sufficient number of MRs. If this reorganization improves distribution resources and abilities, local supply costs can be decreased. Once foreign firms establish sufficient capacity for distribution, they will choose the most profitable distribution channel. This leads to a reduction in foreign supply costs, and only high-cost products are supplied through alliances. This point will be discussed in more detail below. Moreover, there was a trend in M&A among Japanese pharmaceutical wholesalers. Before the 1990s, manufacturers had a close relationship with particular wholesalers.

However, after wholesalers were reorganized, each wholesaler traded with more manufacturers than previously. This also gave foreign firms a chance of creating a distributional channel.

So far, firms have been assumed to conduct Bertrand competition, however an alternative firm interaction should be taken into account: collusion. Two possible collusion schemes are examined: full and partial collusion. I use the inversion formula,  $c = p - \hat{\Omega}^{-1}S$ , to derive marginal costs, where  $\hat{\Omega}$  is a price derivative matrix under a collusion scheme. I found that marginal costs and margins are positive, so I cannot reject collusion schemes by simply deriving marginal costs, because the results do not conflict with profit maximizing behavior. In the following section, by regressing the marginal costs on product characteristics, I test whether collusion schemes are rejected in favor of competition schemes.

While in Figure 1 the two percentiles of the marginal costs do not overlap, Figure 2 shows that these intervals of markups do overlap. No significant difference between markups is found. The average markup rates are quite low: 0.5 to 2 percent. This may be because the antibiotics market in Japan is a mature market. The regulated retail price will be set with a sufficiently large margin when a new drug is launched. However, many antibiotics were launched before 2000, so the regulated retail prices were reset at a low level. For example, Cefdinir, a cephalosporin antibiotic, was launched in 1991. The first regulated per-patient-day price was approximately 500 yen. However, in the 2000s, the prices were reset to approximately 200 yen. Production and distribution costs have not drastically decreased; therefore, the average margins are low in my sample period.

== Figure 2 Here ==

Similar results are obtained by considering the partial collusion case (Figure 3). Under the partial collusion case, these markups do not differ from the competition case. Even if firms forming alliances cooperate, market distortions may not be severe compared with the competition case. However, if we assume that all firms in this market collude, then margins are significantly higher than other competition

and collusion cases (Figure 4). Therefore, it may be crucial to identify what kind of firm conduct prevails to assess efficiency in this market.

== Figure 3 Here ==

== Figure 4 Here ==

## 5 Alliance and Collusion Schemes

In this section, first, I use a positive royalty rate and derive markups consistent with it. Hence, my numerical examination detects marginal costs and price–cost margins consistent with the preset royalty rate by using observed price, quantity, and demand structures. Then, I test which alliance scheme, only fixed fee or positive royalty rate, or collusion scheme, full collusion or partial collusion, is most consistent with the data. Because I have limited data, my approach is similar to that of Villas-Boas (2007) and Bonnet and Dubois (2008), who examined retail pricing behaviors without wholesale price data. The comparison between collusion and competition was conducted by Goldberg and Verboven (2001) for European car markets.

### 5.1 Royalty Rates

As described in Section 3, when the royalty rate is  $\lambda$ , the payment from licensee to licensor is  $\sum_m \lambda p_m S_m M$ . Firms set their prices with these payments taken into account. The FOCs with respect to prices are given by maximizing the profit function (1). This allowed me to solve for the marginal costs,  $c$ , then derive price–cost margins,  $p - c$ .

I employ a numerical examination by using the positive value of  $\lambda$ . Figure 5 shows the average market margins when  $\lambda = 0.1$ . By comparing these with the results when  $\lambda = 0$  in the previous section, my numerical examination demonstrates that high royalty rates are consistent with high markups.

Margins are significantly different between alliances and firms' own products. This result suggests that the presence of royalty rates affects pricing behavior. Therefore, my results confirm a public policy trade-off between benefits from entry through alliances and costs of high markups when high royalty rates are imposed.

== Figure 5 Here ==

The remaining issue concerns whether high distortion is in fact realized in this market. High distortion can also be caused by collusion, as shown in Figure 4. In the following, I compare competition and collusion cases by fitting the functions to marginal costs, and test which case is more consistent with the data.

## 5.2 Tests on Alliance and Collusion Schemes

Because there is no information about competition mode in this market, it is not clear whether competition or collusion prevails in this market. If data on actual margins were available, tests of competition or collusion could be conducted by comparing actual margins and theoretical margins derived from competition schemes as in Ohashi (2002). In this study, however, because of this limited information, I use a similar method to that of Villas-Boas (2007), where wholesale price schemes (e.g., linear or nonlinear pricing) are investigated without wholesale price data. Instead of pricing schemes, I investigate whether a competition scheme is more consistent with the data than a collusion model. A similar exercise is performed by Goldberg and Verboven (2001) in examining whether the UK car market is colluded or not by comparing pricing equations under competition and collusion schemes.

Using implied marginal costs enables us to examine the fit of each model to the data. The log of the marginal cost of product  $i$  is assumed to be linear in product characteristics:

$$\ln c_i = w_i \gamma + \epsilon_i,$$

where  $c_i$  is marginal costs,  $w_i$  is the matrix of covariates, and  $\gamma$  is the parameter vector. To take into account the kinds of elements that affect marginal costs, I use product characteristics as covariates. In addition, the index of foreign subsidiary reorganizations is incorporated to capture the effects of reorganization on supply costs. This variable takes a value of 1 after subsidiary reorganization.

By considering competition and setting several royalty rates ( $\lambda = 0, 0.05, 0.1$  and  $0.5$ ) and considering collusion schemes, the corresponding implied marginal costs are calculated. I regress marginal cost on the covariates by maximum likelihood and consider which estimate fits the data better. As mentioned above, interpretations of each parameter are not shown, but overall fit of the function is examined. Because the data generating process of marginal costs is different between alliance and collusion models, I use a nonnested model selection test statistic. Table 4 reports results for the Vuong test (Vuong (1989)). The Vuong test is based on a likelihood ratio test and makes a pairwise comparison between two models. This test statistic has standard normal distribution. The test statistics are in each cell of Table 4 and the chosen model is denoted in parentheses. If this statistic is positive and large, the model in the column is chosen, and if negative and large, the model in the row is chosen.

Table 4 shows that zero royalty rates are not rejected for all collusion schemes. The results for positive royalty rates, 5 and 10 percent, are not decisive. On the other hand, 50 percent royalty rates are rejected by the partial collusion scheme. This implies that actual average royalty rates are less than 50 percent. Therefore, my results suggest that my sample alliances do not have high royalty rates, and may even have only fixed fees. While this result is obtained under the condition that all alliances have a common value of  $\lambda$ , it can be interpreted that on average low royalty rates are imposed, and therefore there may be a small margin in this market.

Table 4 also reports the results of a comparison among the collusion schemes. The full collusion scheme is rejected by the partial collusion case. The partial collusion case is that in which alliance firms

are considered a single firm. The partial collusion case is not rejected by competition schemes, except for the zero royalty case. Even if there are only fixed fees, alliance agreements can increase the threat of punishment through an alliance in a repeated game situation (Lin (1996)). However, my test results indicate that in this market, pharmaceutical firms do not collude using fixed-fee contracts. These test results are robust under the different specifications of the function, such as including time dummies.

== Table 4 ==

Theoretically, the weak anticompetitive effects are due to product characteristics and market segmentation. If products are a close substitute and firms compete in price, royalty rates will be high, and thus prices charged will also be high (Fauli-Oller and Sandonis (2002)). On the other hand, if products are not substituted and markets are segmented, the industry gains from supplying new products may dominate the loss in industry profits (Tirole (1988)). Therefore, empirical results suggest that in the Japanese antibiotics market, independently developed drugs and alliance drugs are well differentiated. Moreover, because cross-border alliances prevail in this market, the majority of originator firms forming alliances are foreign firms, so the risk of cannibalization of their own product sales is low. There may be less need to manipulate marginal costs to maintain competitive advantage in the product market, or to collude.

## **6 Conclusion**

This study has examined the relationship between cross-border alliances and market competition. By estimating demand for antibiotics and considering Bertrand competition, I derive markups and marginal costs. My estimations show that the marginal costs of products through alliances and products that foreign firms supply by themselves are significantly different. This suggests systematic differences in supply modes. However, foreign supply costs are decreased when foreign subsidiary reorganizations

and regulatory reforms occur. Furthermore, with royalty payments, high royalty rates are consistent with high markups. Because my empirical tests suggest that royalty rates can be zero, distortions caused by high markups may not be severe in this market. Empirical tests also indicate that firms are not colluded. Therefore, the anticompetitive effects of alliance are not found in this market.

Because trade barriers, such as tariffs, tend to be reduced because of WTO or free trade agreements, nontariff barriers become significant. This paper deals with the issue that local supply costs may be a burden for foreign firms. As our results suggest, foreign supply costs may have decreased when domestic regulatory reforms came into effect. Thus, liberalizing the distribution sector will be an important trade negotiation issue. Our results indicate that further liberalizations may improve efficiency.

With respect to supply costs of domestic and foreign firms, my results demonstrate that while foreign supply costs decreased, domestic supply costs increased slightly. Even though Japanese pharmaceutical companies engaged in M&A activities in my sample period, synergy effects were not realized. Because my empirical tests imply that collusion does not seem to occur in this market, a pro-M&A or pro-alliance policy may be required.

My empirical study attempts to examine the significance of the effects of alliance on competition, and thus shed light on the quantitative effect of alliance on firm pricing behavior, which is a major issue in the antitrust literature. Because this study is, to my knowledge, the first attempt to examine the relationship between alliance and market competition, there are still many important issues remaining. In this paper, royalty rates are treated as given and empirically tested. Royalty rates reflect bargaining powers of originators and partners, and therefore affect entry patterns. It would be interesting to investigate how royalty rates are determined. This requires a model for bargaining, and is thus a topic for future research.

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Table 1: Summary Statistics

Variables	Average	Std	Min	Max
first year	1989.241	5.39523	1971	2002
half-life	3.378	6.917	0.5	62
indication	42.714	13.179	14	69
contraindication	1.349	1.558	0	5
foreign	0.228	0.421	0	1
per-day price (oral)	561.233	904.466	108	3800
per-day price (injection)	3828.15	3678.806	1290	15200
share	0.012	0.017	0.0001	0.076
num. of obs = 241				
num. of oral = 171				
Year 1999–2006				

Table 2: Summary Statistics II

Price	Average	Std	Min	Max	Obs
Own (oral)	526.786	816	108	3760	132
Own (injectable)	3930.071	3866.8	1290	15200	63
Domestic Own (oral)	513.097	801.289	108	3760	101
Domestic Own (injectable)	4810.474	4715.244	1290	15200	39
Alliance (oral)	677.822	1160.271	144.9	3800	39
Alliance (injectable)	2910.857	201.077	2646	3200	7
Foreign Alliance (oral)	698.49	1206.604	144.9	3800	36
Foreign Alliance (injectable)	2910.857	201.077	2646	3200	7
Foreign Supply (oral)	571.387	874.442	151.2	3255	31
Foreign Supply (injectable)	2499.417	349.307	1860	3200	24
Weighted Price					
Own (oral)	4.866	5.15	0.157	19.647	132
Own (injectable)	4.506	2.53	1.018	9.119	63
Domestic Own (oral)	5.814	5.463	0.235	19.647	101
Domestic Own (injectable)	5.644	2.34	2.583	9.119	39
Alliance (oral)	2.206	1.857	0.157	6.536	39
Alliance (injectable)	1.297	0.492	0.313	1.879	7
Foreign Alliance (oral)	2.346	1.866	0.157	6.536	36
Foreign Alliance (injectable)	1.297	0.492	0.313	1.879	7
Foreign Supply (oral)	1.779	1.862	0.157	7.436	31
Foreign Supply (injectable)	2.656	1.562	1.018	5.871	24

Table 3: Demand Estimation

	OLS	IV	OLS	IV	OLS	IV
price (in 100 yen)	-0.032c (0.003)	-0.044c (0.006)	-0.031c (0.002)	-0.064c (0.014)	-0.031c (0.002)	-0.071c (0.015)
Drug Age	-0.069c (0.016)	-0.078c (0.014)	-0.084c (0.012)	-0.104c (0.014)	-0.085c (0.012)	-0.113c (0.015)
Half Life	0.013 (0.012)	0.006 (0.008)	0.003 (0.009)	-0.015 (0.013)	0.012 (0.010)	-0.043 (0.048)
Indication	-0.030c (0.007)	-0.039c (0.009)	-0.038c (0.005)	-0.061c (0.009)	-0.038c (0.005)	-0.071c (0.013)
Contraindication	-0.038 (0.056)	-0.052 (0.054)	-0.018 (0.041)	-0.060 (0.041)	-0.021 (0.041)	-0.054 (0.046)
Foreign supply	-1.157c (0.186)	-1.154c (0.155)	-0.878c (0.139)	-0.917c (0.232)	-0.856c (0.139)	-0.848c (0.253)
year2000dummy	-0.046 (0.314)	-0.034 (0.286)	0.001 (0.231)	0.027 (0.301)	-0.046 (0.232)	0.159 (0.466)
year2001dummy	0.103 (0.316)	0.132 (0.287)	0.247 (0.233)	0.302 (0.309)	0.226 (0.232)	0.427 (0.408)
year2002dummy	0.008 (0.321)	0.045 (0.304)	0.137 (0.237)	0.215 (0.312)	0.017 (0.248)	0.561 (0.706)
year2003dummy	-0.031 (0.321)	0.020 (0.308)	0.224 (0.238)	0.321 (0.333)	0.156 (0.241)	0.614 (0.567)
year2004dummy	-0.201 (0.313)	-0.151 (0.317)	0.169 (0.232)	0.243 (0.367)	0.066 (0.240)	0.663 (0.759)
year2005dummy	-0.030 (0.317)	0.033 (0.324)	0.388 (0.236)	0.490 (0.386)	0.339 (0.237)	0.815 (0.590)
year2006dummy	0.031 (0.324)	0.106 (0.331)	0.442a (0.240)	0.578 (0.403)	0.303 (0.255)	1.113 (0.940)
Within share			0.650c (0.047)	0.541a (0.304)	0.653c (0.047)	0.807b (0.329)
Markup rate					-0.762 (0.476)	1.711 (3.416)
Constant	-17.889c (0.445)	-17.195c (0.510)	-15.733c (0.363)	-14.206c (1.346)	-15.454c (0.402)	-13.482c (1.410)
R-squared	0.400	0.359	0.674	0.366	0.676	0.158
Chi square p value		0.09		0.17		0.104
N	241	241	241	241	241	241

The numbers in parentheses are standard errors, and a, b, and c indicate statistical significance at the 1, 5, and 10 percent levels, respectively.

Table 4: Vuong Test

	Partial Col	Full Col
$\lambda = 0$	3.519 (0)	48.146 (0)
$\lambda = 0.05$	-1.495 (?)	21.334 (0.05)
$\lambda = 0.1$	-1.65 (?)	11.142 (0.1)
$\lambda = 0.5$	-3.268 (Par Col)	-0.949 (?)
Partial Col		45.622 (Par Col)

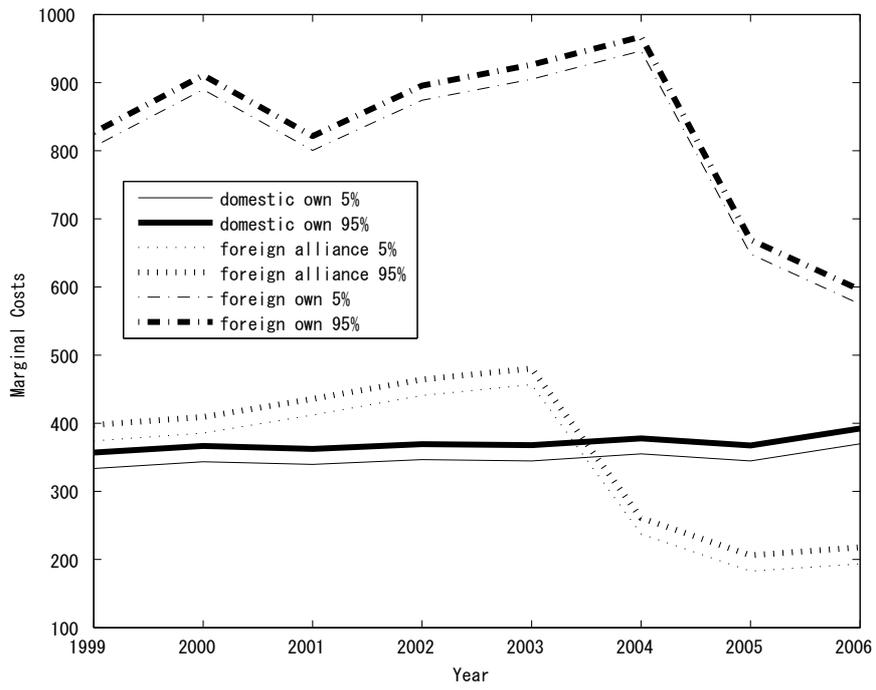


Figure 1: Confidence Intervals of Marginal Costs in Each Year

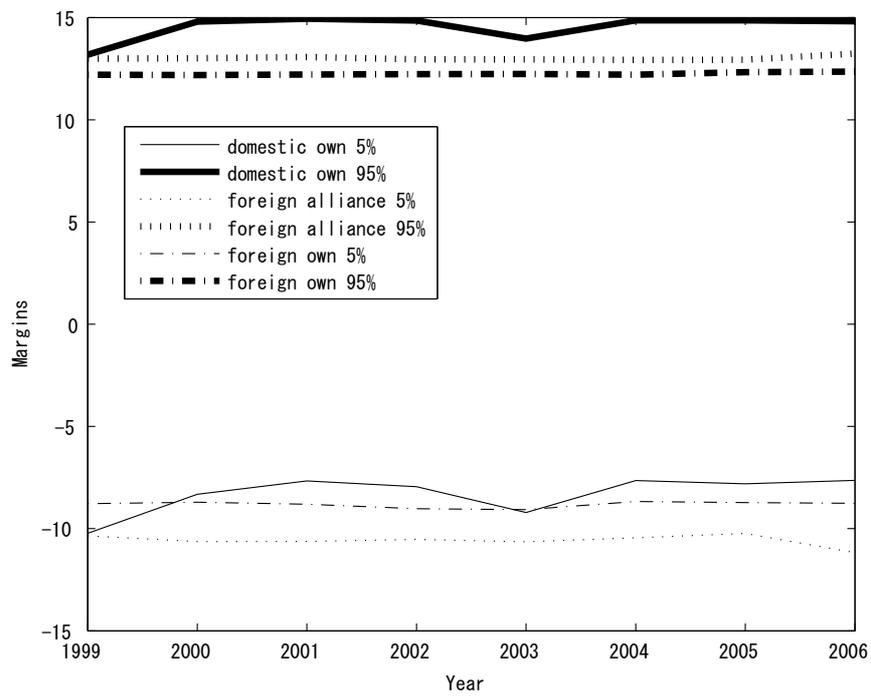


Figure 2: Confidence Intervals of Margins in Each Year

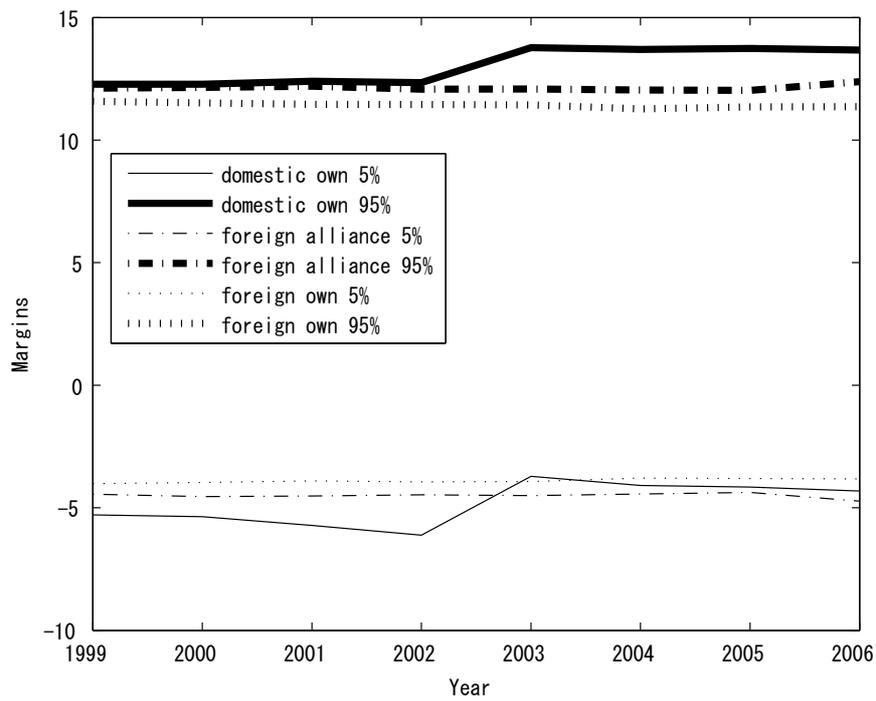


Figure 3: Confidence Intervals of Margins in Each Year (Partial Collusion)

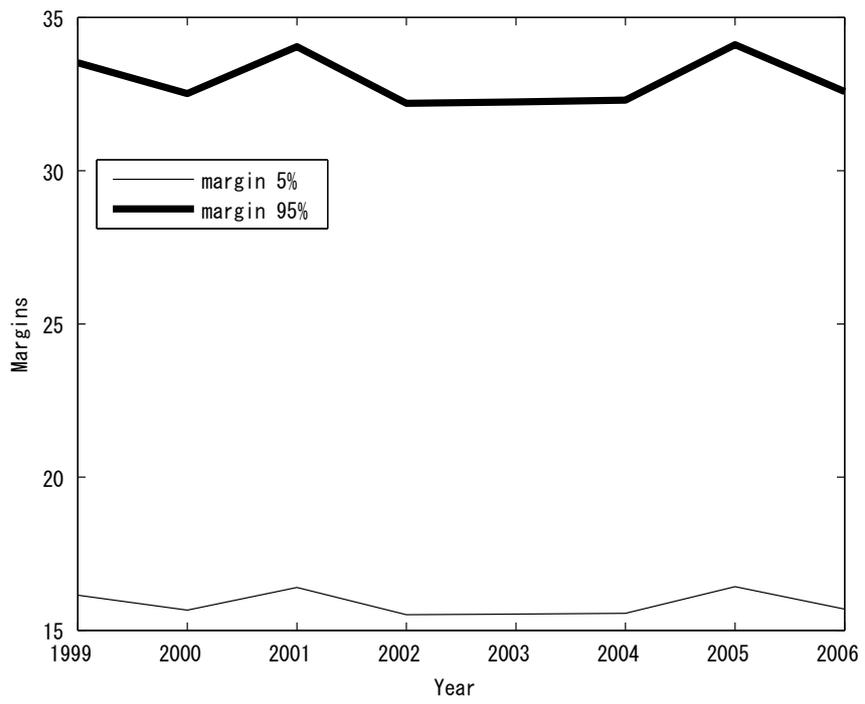


Figure 4: Confidence Intervals of Margins in Each Year (Full Collusion)

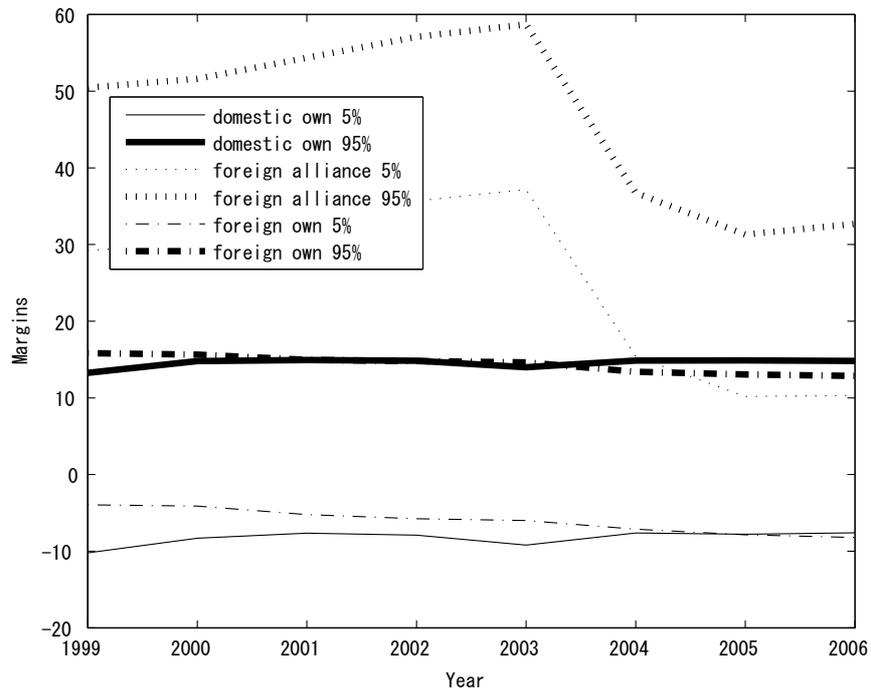


Figure 5: Margins when  $\lambda = 0.1$  in Each Year