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The Effect of Large-Scale Retailers on Price Level: Evidence from Japanese data for 1977-1992

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The Effect of Large-Scale Retailers on Price Level: Evidence from Japanese data for 1977-1992[1]

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Abstract

Since its enactment in 1974 until its easing in the 1990s, the Large-Scale Retail Store Law (*Daikibo Kouri Tenpo Ho*) strictly regulated the entry of large-scale retailers in cities in Japan to protect local small and medium incumbent stores. This paper investigates the effect of large-scale retailers on the price level in Japan using city-level panel data from 1977 to 1992, the period when the Large-Scale Retail Store Law exercised strong entry restrictions. Using fixed effects estimation and instrumental variable estimation, we find that the presence of large-scale retailers, measured by their floor area relative to that of all of the retailers, has a negative effect on the price index of agricultural products, mass-produced food products, textiles, and durable goods. The estimation results suggest that a 10% increase in the relative floor area of large-scale retailers reduces the price level by around 0.3%-1.3%.

Keywords: Large-Scale Retail Store Law, Price level, Panel data, Price index, Japan.
JEL classification: L5

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1 Introduction

Since its enactment in 1974, the Large-Scale Retail Store Law (*Daikibo Kouri Tenpo Ho*) had strictly regulated the entry of large-scale retailers in order to protect local small and medium incumbent stores. Under the Large-Scale Retail Store Law, when a large-scale retailer (LSR) wanted to open a store with a floor area of 500 square meters or more, it had to obtain an approval from the Ministry of International Trade and Industry (METI), which was based on recommendations from a local committee that includes representatives of local incumbent merchants. The local committee could make "adjustments" to the business plan of the applicant, for example, to reduce its floor space, to shorten its operating hours, and to increase the number of non-business days. Furthermore, local authorities were allowed to impose additional restrictions on the potential entrant LSR.

Since the 1990s, the Large-Scale Retail Store Law has been amended to ease the opening of new LSRs and to stop excessive intervention by local business committees. The law was revised in 1992, and, among other things, (i) the period between a business permit application and its approval was shortened to one year, (ii) the local committee was abolished, and (iii) excessive measures by local authorities were restrained. In the 2000s, the Large-Scale Retail Store Law was abolished and replaced with the Large-Scale Retail Store Location Law (*Daikibo Kouri Tenpo Ricchi Ho*).

As described above, the Large-Scale Retail Store Law exercised strict entry restrictions on LSRs from the mid-1970s to the 1990s. In view of consumer welfare, it is of great interest to examine the effect of these restrictions on the local consumer price level in Japan, or, in more quantifiable terms, what was the impact of the presence of LSRs on the local retail price in Japan. Very few studies, however, investigate the effect of the presence of LSRs on the retail price in Japan. This paper aims to fill this gap by examining the price effect of LSRs using a city-level panel dataset that spans the period 1977–1992, when the Large-Scale Retail Store Law exercised strong entry restrictions. An attractive feature of city-level price data is that it allows us to construct a panel dataset for a sufficiently long time period so that there is sufficient variation in the presence of LSRs. While we use only an aggregated dataset for both price level and LSRs, the lack of detailed geographic characteristics in our aggregated data should be at least partly offset by the panel structure of our dataset. We control the city-specific effect by fixed effects estimation, and the endogeneity of the entry of

LSRs by the instrumental variable method. The estimation results suggest a negative effect of LSRs on the price of agricultural products, mass-produced food products, textiles, and durable goods. The estimation results imply that a 10% increase in the floor area of LSRs relative to the total retail floor area reduces the price level by around 0.3–1.3%.

The remainder of this paper is organized as follows. Section 2 reviews the existing studies on the effect of LSRs on local pricing in Japan. Section 3 describes the dataset we use in this study. Section 4 discusses the empirical methodology. Section 5 reports the estimation results, and Section 6 concludes.

2 Literature review

Studies on the effect of LSRs on local pricing in Japan are sparse. Yamashita et al. (1992) analyze the effect of the presence of LSRs on local retail price using detailed city-level and item-level *kohyo* price data from the National Survey of Prices (*Zenkoku Bukka Toukei Tyosa Houkoku*) and from *kohyo* store-level LSR data from the Census of Commerce (*Shogyo Toukei Hyou*) that spanned 681 cities in 1977 and 713 cities in 1987. Their price data cover 55 perishable foods, 117 mass-produced food products, and 41 consumer electronics. They use the sales share of LSRs in each city as their main explanatory variable, and run separate cross-sectional regressions for each of these products and for years 1977 and 1987.

Yamashita et al. (1992) find that when the dependent variable is the price of perishable foods, the coefficient of the sales share of LSRs is significantly different from 0 in about 50% of the items, of which about 80% are positive and about 20% are negative. Further, the average value of the significantly positive coefficient estimates is around 25, whereas the average value of the significantly negative coefficient estimates is around -26. It is difficult to draw an implication for aggregate price level from their result due to such problems as multiple hypotheses testing and weights for aggregation. Taking a simple average of these significant estimates gives $0.8 \times 25 + 0.2 \times (-26) = 14.8$, which suggests that a 1% percent increase in the sales share of LSRs will *increase* the price level of perishable foods by 0.148%. When the dependent variable is the price of mass-produced food products, the coefficient of LSR sales share is significantly different from 0 in about 50% of the items, and the simple average of the estimates is about -6. When the dependent variable is the price of consumer electronics, the simple average of significant estimates is about −8.1. Therefore, for mass-produced food products and consumer electronics, the presence of LSRs appears to have a negative effect on their price.

Motivated by the high consumer price in Shizuoka City, Yamashita et al. (2000) investigate the relationship between retail price and LSRs using city-level price index data from 1992, *Zenkoku Bukka Toukei Tyosa*, and LSR data from Toyo Keizai, *Chiiki Keizai Souran*, and Shogyokai, *Super Market Meikan*. They compare 10 capital cities with similar land prices and find that the price index is negatively correlated with relative sales floor area of LSRs to total retail store floor area. They also find that the market concentration in retailing, measured by the four-firm concentration ratio in retailing, has a negative but insignificant effect on the price index.

Abe and Kawaguchi (2010) use scanner price data that span from 2000 to 2007 and examine the effect of new store openings by Ito-Yokado and Eion on incumbents' pricing of national-brand products. Their price data on incumbent stores are based on weekly scanner data that precisely record the name of each product, time of sale, price, number of units sold, and street address of the store. The street address is used to select the stores affected by new supermarket openings (treatment group) and the stores located nearby but not directly affected by new store openings (control group). Given the detailed geographic information in their dataset, Abe and Kawaguchi (2010) can control for unobserved market heterogeneity and use the differences-in-differences method.

Abe and Kawaguchi (2010) find that stores in the treatment group reduce the prices of national brand curry paste, bottled tea, and instant noodles by 0.4–3.1% after the opening of a new supermarket compared with stores in the control group. Very limited or no effects was found on the prices of detergent and toothpaste. Further, Abe and Kawaguchi (2010) find that the magnitude of incumbents' price reduction depends on the competitiveness of the local market (namely, the number of incumbent supermarkets) and the characteristics of the entrants (namely, whether the entrant is a general merchandise store or a supermarket).

Regarding the economic effect of LSRs, Matsuura and Sugano (2009) use data from 1997 and 2004 and compare the performance of small and medium enterprises at the regional level between regions with/without LSR entry. They find that LSR entry does not have any negative effect on the performance of small and medium enterprises.

3 Data

We use city-level data from cities that had a population of more than 100,000 as of October 1, 1976. The data on price are collected from the National Survey of Prices (*Zenkoku Bukka Toukei Tyosa Houkoku*) published by the Statistics Bureau of Management and Coordination Agency, in 1977, 1982, 1987, and 1992 (Management and Coordination Agency, 1977, 1982, 1987, 1992). As a measure of price level, we use the regional difference index of consumer prices, where the price index of Tokyo ku-area is set to 100. Regarding the choice of commodities, we use four subgroups from Special Group: Commodities (*Tokusyu-Bunrui, Syohin*); these are agricultural and aquatic products, mass-produced food products, textiles, and durable goods. We choose these subgroups because these four groups cover most commodities.

The data on the floor area of LSRs are constructed using the information in *Comprehensive List of Large Scale Retailers in Japan 1992* (*Zenkoku Ogata Kouriten Souran 1992*) published by Toyo Keizai Shinposha (Toyo Keizai Shinposha, 1992). The *Comprehensive List of Large Scale Retailers* contains information on street address, floor area, store opening date, and store opening application date for all LSRs that existed in 1992. From the data on store opening date, we can calculate the floor area of all LSRs in each city in each year and month.

The National Survey of Prices was conducted in May 1977, November 1982, November 1987, and November 1992. To match the timing of the price survey, the floor area of LSRs in 1977, 1982, 1987, and 1992 is defined as the floor area of LSRs that opened before (not including) July 1977, January 1983, January 1988, and January 1993, respectively. The floor area of LSRs in each city is converted to the relative floor area of LSRs using the data on the total floor area of all retailers in each city from the Census of Commerce (*Shogyo Toukei Hyou*) published by METI in 1979, 1982, 1988, and 1991, respectively (Ministry of Trade and Industry, 1979, 1982, 1988, 1991). In the regression model, we include a control variable in order to capture the general economic condition that may affect the city's price level. To this end, we use the logarithm of the annual total retail sales in each city from the Census of Commerce.

Our data of the LSR floor area have some shortcomings. First, our floor area data do not include information on LSRs that exited between 1977 and 1992. Therefore, our measure of the LSR floor area overestimates the actual floor area. Alternatively, our estimation results should be interpreted as the effect of fresh openings of LSRs on the price level. Second, for some stores, the store opening application date is recorded but the store opening date is not recorded. For such stores, we use the store opening application date as a proxy for the store opening date. Third, for some cities, both the store opening application date and the store opening date are unavailable for a majority of the stores. This applies to Urawa-shi, Omiya-shi, Chiba-shi, Ogaki-shi, Nagoya-shi, Toyohashi-shi, Ichinomiya-shi, Seto-shi, Toyoda-shi, Okayama-shi, and Kurashiki-shi. We dropped these cities from the sample. Fourth, some cities exhibit an unusual pattern of store openings. For example, Narashino-shi has zero LSR openings after 1974. Again, we dropped such cities from the sample. This leaves us with 158 cities.

4 Empirical methodology

We use the following panel regression model to model our data:

$$y_{it} = \alpha + \beta_1 x_{it} + \beta_2 z_{it} + \delta_1 d_2 + \delta_2 d_3 + \delta_3 d_4 + u_i + \varepsilon_{it}, \quad i = 1, \ldots, n, \quad t = 1, \ldots, 4, \quad (1)$$

where y_{it} is the price index of city i in year t relative to Tokyo ku-area, where the price level in Tokyo ku-area is set to 100. x_{it} is the relative floor area of LSRs, and z_{it} is the logarithm of the annual total sales of the retailers. The time index $t = 1, 2, 3, 4$ corresponds to year 1977, 1982, 1987, and 1992, respectively. d_2 is a dummy variable representing the time effect in 1982. It is 1 when $t = 2$, and 0 otherwise. d_3 and d_4 are dummy variables for 1987 and 1992, respectively. Without the inclusion of time dummy variables, the time effect on price can be confused with the effect of LSRs. u_i is an unobservable city-specific effect. ε_{it} is an idiosyncratic error term that is uncorrelated with u_i and all the other explanatory variables. $(\alpha, \beta_1, \beta_2, \delta_1, \delta_2, \delta_3)$ are the parameters.

As described in Section 3, our dataset of LSRs do not contain information on LSR exits. Further, our data on store opening date may be subject to measurement errors. It is expected that type 1 LSRs (LSRs whose floor area is 1500 square meters or more) exit less frequently than type 2 LSRs, and that the information on type 1 LSRs is less subject to measurement errors. Therefore, we consider another specification that uses the floor area of type 1 LSRs relative to the total retail floor area as a measure of the presence of LSRs. In this specification, the data quality problem of LSR floor area can be mitigated.

Consistent estimation of model parameters requires that the explanatory variables be uncorrelated with the error term $u_i + \varepsilon_{it}$. In our dataset, there are two possible sources of endogeneity (correlation). First, the relative floor space of LSRs, x_{it}, may be correlated with u_i. The price level of city i may be affected by some city-specific factor u_i, and such a city-specific factor may be correlated with the number of and/or floor area of LSRs. This type of endogeneity can be controlled by fixed effects estimation. The fixed effects estimation estimates the following within-transformed model:

$$y_{it} - \bar{y}_i = \beta_1(x_{it} - \bar{x}_i) + \beta_2(z_{it} - \bar{z}_i) + \delta_1(d_2 - \bar{d}_2) + \delta_2(d_3 - \bar{d}_3) + \delta_3(d_4 - \bar{d}_4) + \varepsilon_{it} - \bar{\varepsilon}_i,$$

where $\bar{y}_i = (1/4)\sum_{t=1}^{4} y_{it}$ is the time average of y_{it} for the ith city, and $\bar{x}_i$, $\bar{z}_i$, $\bar{d}_j$, and $\bar{\varepsilon}_i$ are defined analogously. The within-transformed model does not have the city effect u_i because it is "washed out" by the transformation. Consequently, the transformed model does not suffer from the endogeneity caused by the correlation between u_i and x_{it}.

We also apply random effects estimation to model (1). Random effects estimation assumes u_i is uncorrelated with (x_{it}, z_{it}) and estimates the parameter $(\alpha, \beta_1, \beta_2, \delta_1, \delta_2, \delta_3)$ by generalized least squares. If u_i is indeed uncorrelated with x_{it} and z_{it}, the random effects estimator is better than the fixed effects estimator because it is more efficient. If u_i is correlated with (x_{it}, z_{it}), however, the random effects estimator is inconsistent. The validity of the random effects estimator (namely, the validity of the uncorrelatedness assumption between u_i and (x_{it}, z_{it})) can be tested empirically. One may test the null hypothesis of the random effects model against the alternative hypothesis of the fixed effects model by applying random effects estimation to the model

$$y_{it} = \alpha + \beta x_{it} + \delta_1 d_2 + \delta_2 d_3 + \delta_3 d_4 + \gamma_1 \bar{x}_i + \gamma_2 \bar{z}_i + u_i + \varepsilon_{it}, \quad i = 1, \ldots, n, \quad t = 1, \ldots, 4,$$

in which the time-averages of x_{it} and z_{it}, $\bar{x}_i$ and $\bar{z}_i$, are added to model (1). The value of $\hat{\gamma}_1$ that is significantly different from 0 implies the rejection of the null hypothesis of the random effects model.

The second source of endogeneity is the possible correlation between x_{it} and ε_{it}. This correlation is true if the entry decision of an LSR in city i at time t is influenced by ε_{it}, which includes an unobservable city-/time-specific factor. We mitigate this endogeneity problem by instrumental variable estimation, where the lagged value of x_{it} is used as the instrument. A similar instrumental variable is used in Basker (2005), who examines the effect of Wal-Mart openings on the pricing of local incumbents using U.S. city-level quarterly price survey panel data. For $t = 2, 3, 4$, we use $x_{i,t-1}$ as the instrument for x_{it}. Because one time period in our dataset is five years, the lagged value of x_{it} is unlikely to be correlated with ε_{it}. For $t = 1$ (year 1977), our instrument for x_{it} is the ratio between the LSR floor area in July 1975 and the total retail floor area in 1974, which is obtained from the Census of Commerce. Instrumenting x_{it} is also expected to mitigate the possible measurement problem in x_{it}.

The standard errors are computed using an estimator that allows general heteroscedasticity and serial correlation in ε_{it} (see Arellano (1987)).

5 Estimation results

We do not report the results from the random effects instrumental variable estimation. This is because, as shown in the following tables, the random effects model is rejected against the fixed effects model in many cases when we do not use instruments, and the coefficient estimates from the random effects model are not economically reasonable in many cases.

Table 1 reports the estimation results when the dependent variable is the price of agricultural and aquatic products. The first panel shows the results from fixed effects estimation. The coefficients of the time effect for 1982 and 1992, d_2 and d_4, are negative and significant at the 5% level. These estimates suggest that the price difference between Tokyo and the other areas decreased between 1974 and 1982 further in 1992. The coefficient of log(sales) is positive as expected. The coefficient of LSRs is negative and significant at the 5% level. The estimated value, -5.94, implies that a 10% increase in the floor area of LSRs relative to the total retail floor area will decrease the price level by 0.594%. According to the Census of Commerce 1994, the relative floor area of LSRs in the 13 largest cities in Japan is 43%. Hence, a 10% increase in the relative LSR floor area corresponds to an around 20% increase in the total LSR floor area.

This estimate would not be too far from recent empirical evidence: Abe and Kawaguchi (2010) find that the opening of a large-scale supermarket lowered the prices of curry paste, bottled tea, instant noodles, and toothpaste sold at neighboring local small and medium stores by 0.4–3.1%, where neighboring stores are defined as those within 1.5 km of the new supermarket. Care needs to be taken in interpreting our results, however, because our estimates include a price reduction both from the lower price of LSRs and from the price cutting by local stores competing with LSRs.

The second panel of Table 1 reports the results from random effects estimation. The coefficient of LSRs is negative, but it is close to 0 and no longer significant, which suggests a bias in the random effects estimator. Indeed, the test of the null hypothesis of the random effects model strongly rejects the null with a p-value smaller than 0.01. The third panel of Table 1 reports the results from the fixed effects instrumental variable estimation. All the coefficient estimates are similar to the fixed effects estimates in the first panel. In particular, the coefficient of LSRs changes little and is still significant at the 5% level.

Table 2 reports the estimation results when the dependent variable is the price of agricultural and aquatic products and the relative floor area of type 1 LSRs is used as a measure of the presence of LSRs. The fixed effects estimate of the coefficient of type 1 LSRs is -4.29, which is slightly smaller in absolute value than the corresponding estimate of LSRs in Table 1. The estimate is significant at the 11% level. As in Table 1, the null hypothesis of random effects model is strongly rejected. The instrumental variable estimate of the coefficient of type 1 LSRs is close to 0 and insignificant at any level. For the other coefficients, instrumental variable and fixed effects estimates are similar to each other, including the standard errors.

Table 3 reports the estimation results when the dependent variable is the price of mass-produced food products. The first panel shows the results from fixed effects estimation. Similar to Tables 1 and 2, the coefficients of d_2 and d_4 are negative and significant at the 5% level in many cases. On the other hand, the coefficient of log(sales) is positive but insignificant. The fixed effects estimate of the coefficient of LSRs is -1.39, but is insignificant even at the 20% level. The second panel shows the random effects estimation results. The coefficient of LSRs is positive and insignificant. This is probably due to the bias in the random effects estimator, because the null of the random effects model is strongly rejected. When the instrumental variable estimator is used, the coefficient of LSRs is estimated to be -7.74 and is significant at the 1% level. The estimated value, -7.74, is similar to the corresponding instrumental variable estimate in Table 1.

Table 4 reports the estimation results when the dependent variable is the price of mass-produced food products and the relative floor area of type 1 LSRs is used as a measure of the presence of LSRs. The fixed effects estimate of the coefficient of type 1 LSRs is -1.88, which is similar to the corresponding estimate of LSRs in Table 3 but has a smaller standard error. The estimate is, however, significant only at the 17% level. Similar to Table 4, the random effects estimate of the coefficient of LSRs is positive and insignificant, and the null of random effects is strongly rejected. The instrumental variable estimate of the coefficient of type 1 LSRs is -8.88, which is similar to the corresponding estimate in Table 3, and is significant at the 1% level. Overall, Tables 3 and 4 suggest a negative effect of LSRs and type 1 LSRs on the price of mass-produced food products, but the estimates are imprecise in some cases. According to the instrumental variable estimates, a 10% increase in the relative floor area of LSRs will decrease the price level of food products by around 0.8%.

Table 5 reports the estimation results when the dependent variable is the price of textiles. In contrast to Tables 1–4, the coefficients of d_2 and d_3 are insignificant, and the coefficient of d_4 is positive and significant at the 5% level. Therefore, the price difference in textiles between Tokyo and the other areas increased between 1974 and 1992. The coefficient of log(sales) is not significant at any level. The fixed effects estimate of the coefficient of LSRs is -0.06 and insignificant at any level. The random effects estimate of the coefficient of LSRs is positive but insignificant even at the 30% level. A comparison between the fixed effects and random effects models is inconclusive, because the null of random effects is not rejected at the 40% level. On the other hand, the instrumental variable estimate of the coefficient of LSRs takes a large negative value, -13.3, and is significant at the 1% level.

Table 6 reports the estimation results when the dependent variable is the price of textiles and the relative floor area of type 1 LSRs is used as a measure of the presence of LSRs. In this case, the coefficient of type 1 LSRs is estimated to be positive by both the fixed effects and random effects estimators, even though neither estimate is significant even at the 20% level. Similar to Table 5, the comparison between the fixed effects and random effects models is inconclusive. The instrumental variable estimate of the coefficient of type 1 LSRs is -12.3 and significant at the 5% level. Overall, when the dependent variable is the price of textiles, the instrumental variable estimator provides a more reasonable estimate than the other estimators, and a 10% increase in the relative floor area of LSRs will decrease the price level by around 1.3%.

Table 7 reports the results when the price of durable goods is used as the dependent variable. Similar to Tables 1–4, the coefficients of d_2 and d_4 are negative and significant in many cases, and the coefficient of log(sales) is positive but insignificant. The fixed effects estimate of the coefficient of LSRs is -2.77 and significant at the 5% level. The random effects estimate of the coefficient of LSRs is close to 0, but the random effects estimate may be biased because the null hypothesis of the random effects model is rejected at the 15% level. In contrast to the cases of the other products, the instrumental variable estimate of the coefficient of LSRs is smaller than the fixed effects estimate.

Table 8 reports the results when the price of durable goods is used as the dependent variable and the relative floor area of type 1 LSRs is used as a measure of the presence of LSRs. The fixed effects estimate of the coefficient of type 1 LSRs is -2.93, which is close to the corresponding estimate of LSRs in Table 7, and is significant at the 5% level. The corresponding instrumental variable estimate is -3.27, which is similar to the fixed effects estimate, and is also significant at the 5% level. Overall, the results in Tables 7 and 8 suggest a negative effect of LSRs on the price of durable goods, where a 10% increase in the relative floor area of LSRs will decrease the price level by around 0.3%.

6 Conclusion

We investigate the effect of LSRs on retail price using a city-level panel dataset on the price level of some aggregated goods and on the floor area of LSRs. We control the city-specific effect by fixed effects estimation and the endogeneity of the entry of LSRs by the instrumental variable method. The estimation results suggest a negative effect of LSRs on the price of agricultural products, mass-produced food products, textiles, and durable goods. A 10% increase in the floor area of LSRs relative to the total retail floor area will reduce the price level by around 0.3–1.2%. The estimated effect is not inconsistent with the recent evidence from a study using store-level scanner data. The regression results also demonstrate the importance of the city-specific fixed effect and the potential bias caused by ignoring the correlation between the fixed effect and the dependent variable. One drawback of this study is the lack of exit data of LSRs. An investigation using a more detailed data on prices and the floor area of LSRs is an important future topic.

The expansion of LSRs may enhance consumer welfare via a reduction in retail price. On the other hand, a reduction in retail price may reduce the welfare of local small and medium incumbent stores by reducing their profitability. Therefore, the overall welfare effect of the expansion of LSRs is not clear and remains a topic for future research.

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Table 1: Estimation results
Dependent variable: Price index of agricultural and aquatic products

		Estimate	Std. Error	t-statistic	p-value
Fixed effects estimation	LSRs	−5.94	2.63	−2.26	0.02
	log(sales)	5.36	1.57	3.41	0.00
	d_2	−1.03	0.49	−2.10	0.04
	d_3	−0.28	0.79	−0.35	0.72
	d_4	−3.79	1.10	−3.45	0.00
	Adjusted $R^2 = 0.147$				
Random effects estimation	LSRs	−0.31	2.20	−0.14	0.89
	(sales)	1.51	0.38	3.97	0.00
	d_2	−0.64	0.30	−2.11	0.04
	d_3	0.79	0.40	1.99	0.05
	d_4	−2.01	0.41	−4.91	0.00
	Adjusted $R^2 = 0.143$				
Test of the null of random effects: p-value = 0.00					
Fixed effects instrumental variable estimation	LSRs	−5.44	2.63	−2.07	0.04
	log(sales)	5.36	1.58	3.39	0.00
	d_2	−1.07	0.49	−2.19	0.03
	d_3	−0.33	0.80	−0.42	0.68
	d_4	−3.85	1.10	−3.50	0.00

Table 2: Estimation results
Dependent variable: Price index of agricultural and aquatic products

		Estimate	Std. Error	t-statistic	p-value
Fixed effects estimation	Type 1 LSRs	−4.29	2.64	−1.62	0.11
	log(sales)	5.54	1.63	3.40	0.00
	d_2	−1.33	0.48	−2.80	0.01
	d_3	−0.68	0.79	−0.86	0.39
	d_4	−4.25	1.12	−3.81	0.00
	Adjusted $R^2 = 0.136$				
Random effects estimation	Type 1 LSR	0.62	2.07	0.30	0.76
	log(sales)	1.48	0.37	4.00	0.00
	d_2	−0.70	0.28	−2.19	0.01
	d_3	0.71	0.38	1.87	0.06
	d_4	−2.09	0.40	−5.23	0.00
	Adjusted $R^2 = 0.143$				
Test of the null of random effects: p-value = 0.00					
Fixed effects instrumental variable estimation	Type 1 LSRs	0.04	2.57	0.02	0.99
	log(sales)	5.27	1.66	3.18	0.00
	d_2	−1.56	0.48	−3.22	0.00
	d_3	−0.93	0.81	−1.15	0.25
	d_4	−4.49	1.14	−3.94	0.00

Table 3: Estimation results
Dependent variable: Price index of mass-produced food products

		Estimate	Std. Error	t-statistic	p-value
Fixed effects estimation	LSRs	−1.39	1.43	−0.97	0.33
	log(sales)	0.25	1.00	0.25	0.80
	d_2	−1.48	0.32	−4.67	0.00
	d_3	−0.46	0.51	−0.89	0.37
	d_4	−1.50	0.72	−2.10	0.04
	Adjusted $R^2 = 0.185$				
Random effects estimation	LSRs	1.14	1.22	0.93	0.35
	log(sales)	0.51	0.19	2.65	0.01
	d_2	−1.77	0.19	−9.32	0.00
	d_3	−0.86	0.24	−3.64	0.00
	d_4	−1.99	0.28	−7.24	0.00
	Adjusted $R^2 = 0.202$				
Test of the null of random effects: p-value = 0.01					
Fixed effects instrumental variable estimation	LSRs	−7.74	1.59	−4.88	0.00
	log(sales)	0.35	1.04	0.34	0.73
	d_2	−0.92	0.33	−2.79	0.01
	d_3	0.24	0.52	0.46	0.64
	d_4	−0.76	0.73	−1.04	0.30

Table 4: Estimation results
Dependent variable: Price index of mass-produced food products

		Estimate	Std. Error	t-statistic	p-value
Fixed effects estimation	Type 1 LSRs	−1.88	1.36	−1.39	0.17
	log(sales)	0.35	1.00	0.35	0.73
	d_2	−1.50	0.30	−4.99	0.00
	d_3	−0.50	0.50	−1.00	0.32
	d_4	−1.56	0.71	−2.21	0.03
	Adjusted $R^2 = 0.187$				
Random effects estimation	Type 1 LSRs	0.55	1.30	0.43	0.67
	log(sales)	0.54	0.19	2.79	0.01
	d_2	−1.71	0.18	−9.69	0.00
	d_3	−0.79	0.22	−3.55	0.00
	d_4	−1.92	0.26	−7.31	0.00
	Adjusted $R^2 = 0.202$				
Test of the null of random effects: p-value = 0.01					
Fixed effects instrumental variable estimation	Type 1 LSRs	−8.88	1.62	−5.46	0.00
	log(sales)	0.78	1.05	0.74	0.46
	d_2	−1.14	0.32	−3.59	0.00
	d_3	−0.09	0.51	−0.17	0.86
	d_4	−1.17	0.74	−1.60	0.11

Table 5: Estimation results
Dependent variable: Price index of textiles

		Estimate	Std. Error	t-statistic	p-value
Fixed effects estimation	LSRs	-0.06	4.10	-0.02	0.99
	log(sales)	-0.06	3.03	-0.02	0.98
	d_2	-0.79	0.91	-0.87	0.39
	d_3	-0.56	1.53	-0.37	0.71
	d_4	4.24	2.10	2.02	0.04
	Adjusted R^2 =0.166				
Random effects estimation	LSRs	2.48	2.73	0.91	0.36
	log(sales)	0.91	0.44	2.05	0.04
	d_2	-1.25	0.49	-2.57	0.01
	d_3	-1.29	0.59	-2.17	0.03
	d_4	3.30	0.71	4.64	0.00
	Adjusted R^2 = 0.202				
Test of the null of random effects: p-value = 0.42					
Fixed effects instrumental variable estimation	LSRs	-13.3	4.61	-2.90	0.00
	log(sales)	0.15	3.03	0.05	0.96
	d_2	0.39	0.92	0.43	0.67
	d_3	0.89	1.52	0.59	0.56
	d_4	5.80	2.10	2.77	0.01

Table 6: Estimation results
Dependent variable: Price index of textiles

		Estimate	Std. Error	t-statistic	p-value
Fixed effects	Type 1 LSRs	1.44	4.42	0.33	0.75
estimation	log(sales)	-0.15	3.07	-0.05	0.96
	d_2	-0.87	0.86	-1.01	0.31
	d_3	-0.65	1.49	-0.44	0.66
	d_4	4.16	2.09	1.99	0.05
	Adjusted R^2 =0.166				
Random effects	Type 1 LSRs	3.44	2.68	1.28	0.20
estimation	log(sales)	0.90	0.43	2.08	0.04
	d_2	-1.25	0.46	-2.70	0.00
	d_3	-1.29	0.57	-2.26	0.02
	d_4	3.28	0.70	4.68	0.00
	Adjusted R^2 =0.182				
Test of the null of random effects: p-value = 0.55					
Fixed effects	Type 1 LSRs	-12.3	5.19	-2.38	0.02
instrumental	log(sales)	0.71	3.07	0.23	0.82
variable	d_2	-0.15	0.87	-0.17	0.86
estimation	d_3	0.15	1.48	0.10	0.92
	d_4	4.92	2.07	2.37	0.02

Table 7: Estimation results
Dependent variable: Price index of durable goods

		Estimate	Std. Error	t-statistic	p-value
Fixed effects estimation	LSRs	-2.77	1.51	-1.84	0.07
	log(sales)	0.36	0.96	0.37	0.71
	d_2	-1.03	0.34	-3.02	0.00
	d_3	-0.23	0.53	-0.43	0.67
	d_4	-0.99	0.71	-1.39	0.16
	Adjusted R^2 =0.09				
Random effects estimation	LSRs	-0.75	1.11	-0.68	0.50
	log(sales)	0.27	0.13	2.10	0.04
	d_2	-1.20	0.22	-5.40	0.00
	d_3	-0.42	0.21	-2.00	0.05
	d_4	-1.19	0.23	-5.27	0.00
	Adjusted R^2 =0.09				
Test of the null of random effects: p-value = 0.15					
Fixed effects instrumental variable estimation	LSRs	-0.98	1.56	-0.63	0.53
	log(sales)	0.33	0.97	0.34	0.73
	d_2	-1.19	0.34	-3.48	0.00
	d_3	-0.42	0.53	-0.80	0.43
	d_4	-1.20	0.72	-1.67	0.09

Table 8: Estimation results
Dependent variable: Price index of durable goods

		Estimate	Std. Error	t-statistic	p-value
Fixed effects estimation	Type 1 LSRs	−2.93	1.47	−2.00	0.05
	log(sales)	0.50	0.95	0.52	0.60
	d_2	−1.12	0.32	−3.47	0.00
	d_3	−0.36	0.51	−0.71	0.48
	d_4	−1.16	0.69	−1.66	0.10
	Adjusted R^2 =0.09				
Random effects estimation	Type 1 LSRs	−0.04	1.12	−0.03	0.97
	log(sales)	0.24	0.12	1.94	0.05
	d_2	−1.26	0.21	−5.89	0.00
	d_3	−0.50	0.20	−2.42	0.02
	d_4	−1.27	0.22	−5.74	0.00
	Adjusted R^2 =0.09				
Test of the null of random effects: p-value = 0.03					
Fixed effects instrumental variable estimation	Type 1 LSRs	−3.27	1.46	−2.24	0.03
	log(sales)	0.52	0.95	0.54	0.59
	d_2	−1.10	0.32	−3.41	0.00
	d_3	−0.34	0.51	−0.67	0.50
	d_4	−1.14	0.69	−1.64	0.10