# "New" econometric evidence for the Baldwin-Richardson (1972)/Miyagiwa (1991) theoretical predictions in government procurement

Shingal, Anirudh

World Trade Institute, University of Bern

21 July 2013

# "New" econometric evidence for the Baldwin-Richardson (1972)/Miyagiwa (1991) theoretical predictions in government procurement

Anirudh Shingal*

July 2013

**Abstract**

Baldwin and Richardson (1972) and Miyagiwa (1991) laid out the conditions under which a home-bias in public procurement is rendered ineffective as a protectionist device. Since then there has been little empirical work on this subject. In this paper, we bridge this gap by building a new dataset from WTO notifications on domestic and foreign purchases by Japanese and Swiss governments at the sector level over 1990-2003 and use it to test the BRM theoretical predictions. Significantly, our empirical results support these theoretical predictions.

**JEL classification:** F10, F13, F14, H57

**Key words:** Government procurement, home-bias, Baldwin-Richardson (1972), GPA, Japan, Switzerland

*Senior Research Fellow, WTI, University of Bern. Address: World Trade Institute, Hallerstrasse 6, CH - 3012, Bern. Tel: +41-31-6313270; Fax: +41-31-6313630; Email: anirudh.shingal@wti.org.

# 1 Introduction

The earliest work on the effects of a home-bias in public procurement has been the formal result derived by Baldwin (1970, 1984) and Baldwin and Richardson (1972). In a partial equilibrium perfectly competitive framework, when imported and domestic goods are perfect substitutes and when government demand for these goods is a fraction of domestic output, then a reduction in imports from the government is compensated by a corresponding increase in the imports of the private sector. Thus the effect of a home-bias in procurement on domestic output and imports is "neutralized" and discriminatory public procurement is rendered ineffective as a protectionist device. Miyagiwa (1991) extended this result to an oligopolistic set-up showing that the "crucial assumption" for the "neutrality proposition" (term coined by Brülhart and Trionfetti, 2004) was perfect substitutability between imported and domestic goods. He also showed the result to be less clear-cut for differentiated goods.

Other theoretical work on this subject includes that by Chen (1995), Trionfetti (1997, 2001), Weichenrieder (2001) and Evenett and Hoekman (2005), but no empirical evidence. The part exception to this is the study by Francois et al. (1997), which compares public and private demand across 85 US industrial sectors and infers the ineffectiveness of home-bias from the "smallness" of public demand. Using EU data, Brülhart and Trionfetti (2001, 2004) show that procurement home-bias matters for industrial location/specialization, but their focus is not the Baldwin-Richardson-Miyagiwa (BRM) theoretical predictions.

In this paper, we bridge the gap in this empirical literature by building a new dataset from WTO notifications on domestic and foreign purchases by Japanese and Swiss[1] governments at the sector level over 1990-2003 and use a different econometric approach to provide a *ceteris paribus* test of the BRM theoretical predictions. Significantly, our empirical results support these predictions.

# 2 Data

Procurement data are assembled from statistical submissions made by the Uruguay Round Agreement on Government Procurement (URGPA) signatories to the Committee on Government Procurement under Article XIX: 5 of the URGPA. Unfortunately, only Canada, the EC, Hong Kong, Japan, Norway and USA has submitted these data regularly since the

[1]Our choice of countries is primarily determined by data availability. Both countries have submitted detailed procurement data sufficiently regularly over 1990-2003 (Japanese procurement data are missing for 1994-1996, the Swiss for 1992) and in a form amenable to econometric analysis.

Uruguay Round[2].

Even amongst the countries that have submitted these data, there are significant differences, both in terms of what is included and how it is included[3], and the need for consistency has thus determined the choice of sample countries for this analysis. For both Japan and Switzerland, we consider all goods[4] procured by central government entities included in Annex 1 of Appendix 1 of the URGPA; the sector descriptions are available in Table 1.

Government procurement rules at the WTO require that only contracts above a certain threshold[5] value be subject to internationally competitive bidding. Article XIX: 5(b) of the URGPA requires submission of data on above-threshold procurement by sector according to the nationality of the winning supplier. This gives us both the value and the number of contracts supplied from abroad by sector.

Looking at these data averages by sector for both countries over 1990-2003 in Table 1, we see that Japanese goods procurement was concentrated in machinery; medical, scientific and photographic equipment; and telecom and electrical equipment (together accounting for 80.4% of total goods procurement), but apart from medical etc. equipment, market access was not high in any of these sectors either by value or number of contracts awarded. In the case of Switzerland, machinery; transport; and medical, scientific and photographic equipment were the three largest sectors accounting for 73.1% of total goods procurement over time and the propensity of Swiss governments to source from abroad was high in these sectors (as well as in agriculture; wood, paper; textiles; and telecom and electrical equipment).

**<Insert Table 1 here>**

We can use private sector import propensities to simulate public sector imports and then use the difference between simulated and actual levels of foreign procurement to derive a measure of the home-bias. Following Shingal (2011), we call this measure the Private-Public Purchase Differential (PPPD)[6]. Significantly, the fitted plot of PPPD against sectoral government

[2]Switzerland has not provided data beyond 2003. A snapshot of country procurement submissions is available at http://wto.org/english/tratop_e/gproc_e/gpstat_e.htm.

[3]For instance, Norway and the US employ a different classification system compared to the EC, Japan and Switzerland which makes it impossible to analyze data at the disaggregated sectoral level for the period under study. Canada provides no information on nationality of winning suppliers. Hong Kong's submissions until very recently have had restricted access.

[4]We exclude services from our analysis as IIP and price data required in our empirical analyses were unavailable for the services sectors.

[5]Thresholds differ depending on the type of procurement and on the level of government making the purchase.

[6]Formally, $PPPD_{kt}$= [{$(M_{kt}$-$V^f_{kt})/Y_{kt}$}*$ATV_{kt}$]-$V^f_{kt}$ where $M_k$= Sectoral total import value, $V^f_k$= Sectoral public import value, $Y_k$= Sectoral output, $ATV_k$= Sectoral above-threshold procurement value.

demand in Figure 1 also suggests that the magnitude of home-bias may have been greater in the large government demand sectors and in line with the Baldwin and Richardson (1972) result for such sectors, the effects of this home-bias are unlikely to be benign.

**<Insert Figure 1 here>**

# 3 Empirical model

The BRM results yield the following testable propositions:

*Proposition 1 (neutrality proposition):* When government demand in a sector as a share of domestic output is low and the procured good is homogeneous, then the sector's total (public and private) imports are independent of the level of foreign procurement.

Formally, for the neutrality proposition to hold, $\frac{\partial M^P}{\partial M^G} = -1$ where $M^G$= public imports and $M^P$= private imports.

Now $\partial M^T = \partial M^G + \partial M^P$so $\frac{\partial M^T}{\partial M^G} = 1 + \frac{\partial M^P}{\partial M^G} = 1 + (-1$ from the neutrality proposition) $= 0$, where $M^T$ = total imports.

Thus, for the neutrality proposition to hold $\frac{\partial M^T}{\partial M^G} = 0$.

*Proposition 2:* In large public demand sectors, a procurement home-bias results in a decline in total imports i.e.$\frac{\partial M^T}{\partial M^G} > 0$.

*Proposition 3 (from Miyagiwa, 1991):* When the procured good is differentiated, the relationship between total and public imports is ambiguous. However, assuming zero conjectural variations, if foreign firm demand is "sufficiently" convex and its total sales "sufficiently" large (Miyagiwa, 1991, pp. 1325), then $\frac{\partial M^T}{\partial M^G} < 0$ i.e. discriminatory public procurement can actually increase total imports.

We use an augmented import demand function (for instance see Warner and Kreinin, 1983) to test the BRM results empirically using the following estimation:

$$M^T_{kt}=\alpha_t + \alpha_k + \Omega_1 IIP_{kt} + \Omega_2 P^M_{kt} + \Omega_3 P^d_{kt} + \Omega_4 M^G_{kt} + \Omega_5 M^G.size_{kt} + \Omega_6 M^G.IIT_{kt} + \Omega_7 size_{kt} + \Omega_8 IIT_{kt} + \vartheta_{kt} \quad \ldots\ldots (1)$$

where "k" denotes the sector, $M^T$ is the volume of total imports, IIP is the index of industrial production that proxies the impact of income on import demand, $P^M$ is the unit value import

price (data on all three variables taken from Nicita and Olarreaga, 2007), $P^d$ is the domestic price level (the Domestic Corporate Goods Price Index from the Bank of Japan and the Producer Price Index from Office Fédéral de la Statistique, Switzerland), $M^G$ proxies the volume of public imports using the number of procurement contracts awarded to foreign firms[7] (compiled from WTO-submitted data), unobserved sector-specific determinants are captured by sector-specific fixed effects ($\alpha_k$) and economy-wide determinants are captured by year fixed effects ($\alpha_t$).

To test for the theoretical predictions in BRM, we interact $M^G$ with "size", which is constructed sectorally as the share of total government demand in domestic output, and with "IIT", which is the Grubel and Lloyd (1971) measure of intra-industry trade[8], and serves as a proxy for the product differentiation within a sector and hence for the (in)substitutability between domestic and imported products.

Formally, $Size_k = \frac{TPV_k^A}{Y_k}$ where $TPV_k^A$= Sectoral measure of total public demand, $Y_k$= Sectoral output. Since the data reporting requirements of the URGPA require data on total procurement to be reported annually but not at the sectoral level, $TPV^A$ is a constructed variable such that $TPV_{kt}^A = TPV_t.\frac{ATV_{kt}}{\sum ATV_{kt}}$ where TPV = Total procurement value, ATV = Above-threshold procurement value. Thus, "$size_k$" $\approx 0$ would denote sectors where public demand was a fraction of domestic output.

The goods in the Japanese and Swiss procurement data are far more aggregated to enable their classification as homogeneous and differentiated on the basis of the Rauch (1999) classification. However, the IIT literature suggests that intra-industry trade is associated with product differentiation and *a la* Yang (1997) we therefore use the extent of IIT in a sector to proxy the extent of product differentiation in that sector. Thus, $IIT_k = 0$ would denote sectors with homogeneous goods implying perfect substitutability.

A priori, we expect $\Omega_1$ and $\Omega_3$ to be positive and $\Omega_2$ to be negative. A statistically significant estimate of $\Omega_4 \approx 0$ or estimated $\Omega_4$ statistically indifferent from zero would validate the neutrality proposition, with size and IIT as additional controls. With "size" as an additional control variable, a positive estimate of $(\Omega_4 + \Omega_5)$ significantly different from zero would validate Proposition 2. With "IIT" as an additional control, in line with Proposition 3, the expected sign of $\Omega_6$ is ambiguous. But a statistically significant estimate of $(\Omega_4 + \Omega_6) \approx 0$ or estimated $(\Omega_4 + \Omega_6)$ statistically indifferent from zero would support the neutrality

[7] There are no data on the volume of public imports in the WTO submissions, but since we are dealing with high value above-threshold procurement, an increase in the number of contracts awarded to foreign firms must on average imply an increase in the volume of public imports.

[8] Formally, $IIT_k = 1 - [\text{abs}(X_k\text{-}M_k)/(X_k + M_k)]$ where $X_k$ = Sectoral export value, $M_k$ = Sectoral import value.

proposition for differentiated goods. Finally, with both "size" and "IIT" in a fully-specified equation (1), the overall estimate of $(\Omega_4+\Omega_5+\Omega_6)$ would depend on the relative magnitudes of $\Omega_5$ and $\Omega_6$ as well as the sign of the latter.

We found the dependent variable in equation (1) to be characterized by over-dispersion[9] which rendered a log-linear OLS estimation biased. Given the scale-dependence of the negative binomial pseudo-maximum likelihood (PML) estimator (Bosquet and Boulhol, 2010), the Poisson-PML (PPML) was our preferred estimator. The adequacy of the PPML was also successfully tested using the Gauss Newton regression in Silva and Tenreyro (2006)[10].

# 4 Results

Table 2 reports the results from estimating equation (1) on Japanese and Swiss data pooled together. Given that the Japanese economy has a larger government and is also less open than Switzerland[11], we estimate equation (1) on the pooled sample with country fixed effects[12]. While we focus on results from the PPML estimation reported in columns VI-X, for the sake of comparison, we also report results from a standard log-linear OLS estimation in columns I-V. As a robustness check, we also replaced $TPV_k^A$ with $ATV_k$ in the definition of our "size" variable but found these (unreported) results to be robust to this change.

From Table 2 we see that the estimate of $\Omega_4$ in specifications VIII through X is both economically and statistically indifferent from zero, which suggests that the neutrality proposition is validated by both "small" public demand and homogeneous goods sectors.

Estimated $(\Omega_4 + \Omega_5)$ is both economically and statistically different from zero (0.011 and 0.014 in specifications VIII and X, respectively), which supports Proposition 2 and suggests

[9] Over-dispersion in the raw data is due to unobserved heterogeneity (Greene, 1994); the description is used for data where the conditional variance exceeds the conditional mean. Formally this is tested by pitting the null of $Var(y|x) = E(y|x)$ against $Var(y|x) = E(y|x) + \alpha.E(y|x)^2$ where $\alpha > 0$ suggests over-dispersion. Following Cameron and Trivedi (2005), the null of $\alpha = 0$ against $\alpha > 0$ is the t-test of $\alpha$ obtained by running the auxiliary (no-constant) OLS regression $\frac{\{y-E(y|x)\}^2-y}{E(y|x)} = \alpha.E(y|x) + \epsilon$ where $E(y|x)$ are the fitted values from estimating the Poisson model. The alternative of $\alpha > 0$ failed rejection conclusively (p-value=1).

[10] Formally, $\frac{\varepsilon^2}{\sqrt{E(y|x)}} = \gamma.\sqrt{E(y|x)} + \gamma(\lambda-1).lnE(y|x).\sqrt{E(y|x)} + \xi$ is estimated using OLS, where $\varepsilon = y - E(y|x)$ and $E(y|x) = exp(x\beta)$ and the statistical significance of $\gamma.(\lambda-1)$ is tested using a Eicker-White robust covariance matrix estimator. The null of $\gamma.(\lambda-1) = 0$ (or $\lambda$=1 i.e. evidence for the PPML) failed rejection at the 5% level of significance (p-value = 0.066).

[11] The average share of total government expenditure in GDP over 1990-2003 was almost 50% in Japan versus 37% in Switzerland, while the average share of trade in GDP in these economies was 19.2 and 74.6%, respectively, over the same period.

[12] Given that "sector" definition is consistent across the two countries, we do not include country-and-sector fixed effects. However, we also estimated equation (1) with country, sector and country-and-year fixed effects, but found these (unreported) results to be qualitatively similar.

that the effect of a procurement home-bias on total imports may not be neutralized in large public demand sectors. This finding holds even when we control for product differentiation within a sector, driving the overall non-zero estimate of $(\Omega_4+\Omega_5+\Omega_6)$ in the full-specification X.

Interestingly, estimates of $(\Omega_4 + \Omega_6)$ are found to be both economically and statistically indifferent from zero in both IX and X, which suggests that the neutrality proposition seems to hold for differentiated goods in our data[13]. Following Proposition 3, this last finding suggests that foreign suppliers to these markets may not have faced sufficiently convex demands and that they may have made insufficient total (public and private) sales to these countries in the differentiated sectors or that their conjectural variations may have been positive (Miyagiwa, 1991, pp. 1325).

**<Insert Table 2 here>**

To further examine these results at the country-level, we also estimated equation (1) on Japanese and Swiss data in separate regressions. The results from these, using the PPML estimator, are reported in Table 3. Though we do not find statistically significant evidence for Proposition 2 in these results for either country, we now find statistically significant evidence (though weak in the case of Japan) for the neutrality proposition for both homogeneous and differentiated goods (estimates of $\Omega_4 \approx 0$, $\Omega_4 + \Omega_6 \approx 0$ in columns IV and IX). This said, given the larger number of observations in the pooled sample, the results reported in Table 2 constitute a more robust test of the BRM theoretical predictions.

**<Insert Table 3 here>**

# 5 Conclusion

We provide new econometric evidence on the Baldwin and Richardson (1972)/Miyagiwa (1991) ineffectiveness proposition in public procurement. Our empirical result on the adverse effects of a procurement home-bias in large public demand sectors provides more support to the "preliminary evidence" (Trionfetti, 2000) in this literature on the impact of discriminatory procurement on trade flows and international specialisation. The result is also significant given the current economic stagnation in advanced economies and their well-documented home-bias in public purchase decisions during the recent crisis (Evenett, 2009a,b).

[13]The OLS estimates, on the other hand, refute all four propositions emanating from the BRM predictions.

# References


Baldwin, R. (1970). *Nontariff distortions of international trade*. Brookings institution Washington.

Baldwin, R. (1984). Trade policies in developed countries. *Handbook of international economics*, 1:571–619.

Baldwin, R. and Richardson, J. (1972). Government purchasing policies, other ntbs, and the international monetary crisis. *Obstacles to Trade in the Pacific Area. Ottawa: Carleton School of International Affairs*.

Bosquet, C. and Boulhol, H. (2010). Scale-dependence of the negative binomial pseudo-maximum likelihood estimator. *Documents de travail du Centre d'Economie de la Sorbonne 10092, Université Panthéon-Sorbonne (Paris 1), Centre d'Economie de la Sorbonne*.

Brülhart, M. and Trionfetti, F. (2001). Industrial specialisation and public procurement: Theory and empirical evidence. *Journal of Economic Integration*, 16(1):106–127.

Brülhart, M. and Trionfetti, F. (2004). Public expenditure, international specialisation and agglomeration. *European Economic Review*, 48(4):851–881.

Cameron, A. and Trivedi, P. (2005). Microeconometrics: Methods and applications.

Chen, X. (1995). Directing government procurement as an incentive of production. *Journal of Economic Integration*, pages 130–140.

Evenett, S. (2009a). The emerging contours of crisis-era protectionism. *Broken promises: A G-20 summit report by Global Trade Alert*, pages 15–24.

Evenett, S. (2009b). What can be learned from crisis-era protectionism? an initial assessment. *Business and Politics*, 11(3).

Evenett, S. and Hoekman, B. (2005). Government procurement: market access, transparency, and multilateral trade rules. *European Journal of Political Economy*, 21(1):163–183.

Francois, J., Nelson, D., and Palmeter, D. (1997). Government procurement in the us: A post-uruguay round analysis. *Law and Policy in Public Purchasing. University of Michigan Press, Ann Arbor*, pages 105–124.

Greene, W. (1994). Accounting for excess zeros and sample selection in poisson and negative binomial regression models. *Stern School of Business, New York University, Working Paper 94-10*.

Grubel, H. and Lloyd, P. (1971). The empirical measurement of intra-industry trade. *Economic Record*, 47(4):494–517.

Miyagiwa, K. (1991). Oligopoly and discriminatory government procurement policy. *The American Economic Review*, 81(5):1320–1328.

Nicita, A. and Olarreaga, M. (2007). Trade, production, and protection database, 1976–2004. *The World Bank Economic Review*, 21(1):165–171.

Rauch, J. (1999). Networks versus markets in international trade. *Journal of international Economics*, 48(1):7–35.

Shingal, A. (2011). The wto's agreement on government procurement: Whither market access? *World Trade Review*, 10(4):1–23.

Silva, J. S. and Tenreyro, S. (2006). The log of gravity. *The Review of Economics and Statistics*, 88(4):641–658.

Trionfetti, F. (1997). Public expenditure and economic geography. *Annales d'Economie et de Statistique*, pages 101–120.

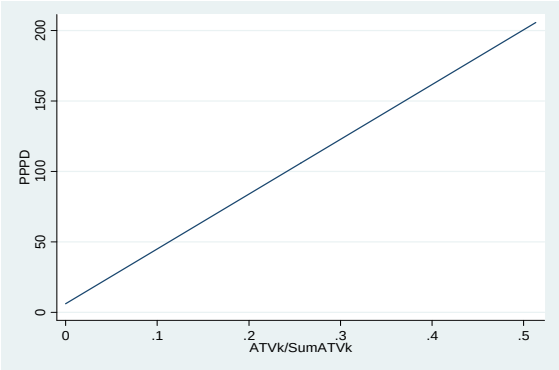
Trionfetti, F. (2000). Discriminatory public procurement and international trade. *The World Economy*, 23(1):57–76.

Trionfetti, F. (2001). Public procurement, market integration, and income inequalities. *Review of International Economics*, 9(1):29–41.

Weichenrieder, A. (2001). Public procurement in the presence of capital taxation. *Regional Science and Urban Economics*, 31(2):339–353.

Yang, J. (1997). Exchange rate pass-through in us manufacturing industries. *Review of Economics and Statistics*, 79(1):95–104.

Figure 1: Home-bias v magnitude of government demand




**Note:** (1) The figure shows fitted values from a linear prediction (2) PPPD is a measure of the sectoral home-bias in public procurement; ATVk/$\sum$ATV is a measure of the relative magnitude of public demand by sector ["ATV" = Above-threshold procurement value] (3) Figure suggests a positive relationship between the sectoral magnitude of home-bias and that of government demand in both countries.

Table 1: Procurement and other data by sector (average 1990-2003)

| **Japan** | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Sector** | **ATV** | **$V^f$** | **$V^f$/ATV** | **ATN** | **$N^f$** | **$N^f$/ATN** | **IIP** | **$M^T$** | **$P^M$** | **$P^d$** | **IIT** | **Size** |
| Agriculture | 0.7 | 0.2 | 25.8% | 1.5 | 0.5 | 29.4% | 99.1 | 14.0 | 1.5 | 97.9 | 0.13 | 0.000 |
| Chemicals, pharma | 215.2 | 72.6 | 33.8% | 1761.1 | 498.3 | 28.3% | 107.9 | 12.0 | 1.8 | 100.6 | 0.82 | 0.003 |
| Plastic, rubber, leather | 18.1 | 0.4 | 2.3% | 33.0 | 0.5 | 1.4% | 97.4 | 1.0 | 7.1 | 101.7 | 0.93 | 0.000 |
| Wood, paper | 198.7 | 8.6 | 4.3% | 262.4 | 10.0 | 3.8% | 92.1 | 13.0 | 0.7 | 101.5 | 0.37 | 0.005 |
| Textiles | 89.5 | 1.6 | 1.7% | 139.3 | 2.5 | 1.8% | 89.2 | 1.9 | 9.6 | 103.5 | 0.52 | 0.004 |
| Stone, ceramic, glass | 3.5 | 0.0 | 1.1% | 9.1 | 0.2 | 2.0% | 91.2 | 3.2 | 0.7 | 101.7 | 0.80 | 0.000 |
| Iron & steel, non-fer metals | 124.1 | 3.4 | 2.8% | 158.2 | 6.5 | 4.1% | 99.7 | 13.0 | 1.1 | 107.1 | 0.85 | 0.002 |
| Machinery | 2438.6 | 199.9 | 8.2% | 1077.3 | 91.4 | 8.5% | 98.5 | 1.1 | 20.2 | 101.4 | 0.48 | 0.018 |
| Telecom & electrical equipment | 881.9 | 58.8 | 6.7% | 626.5 | 41.8 | 6.7% | 119.5 | 15.0 | 21.8 | 111.4 | 0.51 | 0.006 |
| Transport equipment | 364.0 | 56.1 | 15.4% | 351.2 | 31.1 | 8.9% | 107.1 | 0.41 | 20.5 | 101.5 | 0.26 | 0.002 |
| Medical, scientific, photographic equip | 984.6 | 307.6 | 31.2% | 2996.6 | 864.2 | 28.8% | 108.1 | 0.15 | 101.4 | 101.1 | 0.70 | 0.050 |
| Furniture | 37.5 | 0.2 | 0.6% | 100.7 | 0.9 | 0.9% | 106.7 | 0.67 | 4.4 | 99.8 | 0.27 | 0.004 |
| **Switzerland** | | | | | | | | | | | | |
| **Sector** | **ATV** | **$V^f$** | **$V^f$/ATV** | **ATN** | **$N^f$** | **$N^f$/ATN** | **IIP** | **$M^T$** | **$P^M$** | **$P^d$** | **IIT** | **Size** |
| Agriculture | 15.7 | 10.2 | 65.2% | 29.1 | 21.2 | 72.8% | 97.6 | 1.9 | 1.7 | 106.7 | 0.80 | 0.004 |
| Chemicals, pharma | 1.1 | 0.1 | 7.1% | 2.5 | 0.4 | 15.6% | 168.7 | 4.2 | 2.9 | 117.4 | 0.74 | 0.000 |
| Plastic, rubber, leather | 6.2 | 1.7 | 28.0% | 9.5 | 1.4 | 14.6% | 106.1 | 0.31 | 6.7 | 95.0 | 0.82 | 0.003 |
| Wood, paper | 3.2 | 1.7 | 52.6% | 7.8 | 3.6 | 46.5% | 100.6 | 2.5 | 1.0 | 100.9 | 0.80 | 0.003 |
| Textiles | 6.6 | 2.3 | 35.1% | 16.5 | 5.8 | 35.3% | 89.2 | 0.24 | 18.8 | 98.1 | 0.68 | 0.005 |
| Stone, ceramic, glass | 0.3 | 0.0 | 0.0% | 1.9 | 0.0 | 0.0% | 114.9 | 1.9 | 0.6 | 99.2 | 0.71 | 0.000 |
| Iron & steel, non-fer metals | 8.0 | 0.9 | 11.0% | 12.8 | 1.5 | 12.0% | 109.6 | 2.7 | 1.2 | 99.3 | 0.87 | 0.006 |
| Machinery | 112.6 | 82.1 | 72.9% | 165.3 | 118.0 | 71.4% | 109.6 | 0.52 | 20.2 | 96.8 | 0.80 | 0.010 |
| Telecom & electrical equipment | 13.4 | 5.5 | 41.0% | 20.6 | 7.8 | 38.1% | 106.6 | 0.21 | 27.8 | 96.8 | 0.97 | 0.002 |
| Transport equipment | 40.6 | 15.6 | 38.4% | 35.6 | 13.9 | 39.1% | 106.8 | 0.58 | 14.4 | 104.9 | 0.36 | 0.041 |
| Medical, scientific, photographic equip | 22.5 | 12.5 | 55.7% | 33.8 | 21.4 | 63.2% | 107.2 | 0.02 | 83.2 | 98.8 | 0.42 | 0.005 |
| Furniture | 10.2 | 2.9 | 28.2% | 10.0 | 2.1 | 20.8% | 100.1 | 0.31 | 4.8 | 97.1 | 0.48 | 0.007 |

**Source:** WTO (various years); Nicita & Olarreaga (2006); Bank of Japan (various years); Office Fédéral de la Statistique, Switzerland (various years); own calculations

**Note:** (1) "ATV" = Above-threshold procurement by value of contracts; "$V^f$"= Value of contracts awarded to foreign suppliers; "ATN" = Above-threshold procurement by number of contracts; "$N^f$"= Number of contracts awarded to foreign suppliers; rest of the variables are as defined in the paper (2) Units of measurement: ATV, $V^f$(real USD mn); ATN, $N^f$ (units); IIP, $P^M$, $P^d$ (indices); $M^T$ (billion units); IIT, size (ratios)

Table 2: Results: Combined

| | OLS log-linear (dependent variable: $M^T$) | | | | | PPML (dependent variable: $M^T$) | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | **I** | **II** | **III** | **IV** | **V** | **VI** | **VII** | **VIII** | **IX** | **X** |
| IIP | 1.061*** | 1.514*** | 1.723*** | 1.530*** | 1.655*** | 0.007*** | 0.006*** | 0.008*** | 0.006*** | 0.007*** |
| | (0.249) | (0.438) | (0.432) | (0.382) | (0.371) | (0.001) | (0.001) | (0.002) | (0.001) | (0.002) |
| $P^M$ | -1.566*** | -1.600*** | -1.327*** | -1.159*** | -1.042*** | -0.146*** | -0.147*** | -0.137*** | -0.157*** | -0.147*** |
| | (0.160) | (0.208) | (0.236) | (0.189) | (0.211) | (0.018) | (0.018) | (0.021) | (0.018) | (0.022) |
| $P^d$ | 1.732* | 2.934** | 3.529** | 3.127** | 2.953** | 0.010*** | 0.008*** | 0.011*** | 0.006* | 0.007** |
| | (0.774) | (1.082) | (1.262) | (0.953) | (1.096) | (0.002) | (0.002) | (0.002) | (0.003) | (0.003) |
| $M^G$ | | 0.090* | 0.120* | -0.259** | -0.204# | | -0.000 | 0.000 | 0.002 | 0.003 |
| | | (0.039) | (0.054) | (0.092) | (0.113) | | (0.000) | (0.000) | (0.002) | (0.002) |
| Size | | | 9.876** | | 6.397* | | | 5.995** | | 4.678# |
| | | | (2.932) | | (2.629) | | | (2.136) | | (2.489) |
| $M^G$.Size | | | -1.254 | | -0.396 | | | 0.011* | | 0.011# |
| | | | (0.845) | | (0.810) | | | (0.005) | | (0.006) |
| IIT | | | | -2.618*** | -2.513*** | | | | -0.971*** | -0.749*** |
| | | | | (0.417) | (0.503) | | | | (0.154) | (0.200) |
| $M^G$.IIT | | | | 0.626*** | 0.565*** | | | | -0.002 | -0.003 |
| | | | | (0.124) | (0.148) | | | | (0.002) | (0.002) |
| Constant | 10.456** | 2.240 | -1.685 | 2.146 | 2.140 | 21.677*** | 21.957*** | 21.413*** | 22.428*** | 22.053*** |
| | (3.586) | (5.580) | (6.359) | (4.774) | (5.536) | (0.224) | (0.300) | (0.329) | (0.344) | (0.390) |
| # Observations | 258 | 178 | 155 | 178 | 155 | 258 | 258 | 219 | 258 | 219 |
| df_m | 27 | 27 | 29 | 29 | 31 | 27 | 28 | 30 | 30 | 32 |
| r2 | 0.918 | 0.926 | 0.939 | 0.941 | 0.949 | 0.963 | 0.964 | 0.968 | 0.959 | 0.962 |
| Year fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Country fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

**Note:** (1) Legend: # p<.1; * p<.05; ** p<.01; *** p<.001 (2) Standard errors reported in brackets.

Table 3: Results: Japan v Switzerland

| **PPML** | **Japan (dependent variable: $M^T$)** | | | | | **Switzerland (dependent variable: $M^T$)** | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | **I** | **II** | **III** | **IV** | **V** | **VI** | **VII** | **VIII** | **IX** | **X** |
| IIP | 0.004# | 0.004 | 0.003 | 0.000 | 0.000 | -0.001* | -0.001# | 0.000 | -0.001# | 0.000 |
| | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| $P^M$ | -0.074*** | -0.074*** | -0.065** | -0.052** | -0.046** | -0.026*** | -0.030*** | -0.032*** | -0.029*** | -0.032*** |
| | (0.020) | (0.020) | (0.020) | (0.016) | (0.016) | (0.007) | (0.008) | (0.007) | (0.007) | (0.007) |
| $P^d$ | 0.003 | 0.003 | 0.003 | 0.012** | 0.011** | -0.003* | -0.002# | -0.005*** | -0.002* | -0.005** |
| | (0.003) | (0.003) | (0.003) | (0.004) | (0.004) | (0.001) | (0.001) | (0.002) | (0.001) | (0.002) |
| $M^G$ | | 0.000 | -0.000 | 0.004# | 0.003 | | 0.001 | 0.001 | -0.007** | 0.004 |
| | | (0.000) | (0.000) | (0.002) | (0.002) | | (0.001) | (0.001) | (0.002) | (0.008) |
| Size | | | 19.017* | | 12.345# | | | -0.298 | | -0.310 |
| | | | (7.821) | | (6.916) | | | (0.341) | | (0.327) |
| $M^G$.Size | | | -0.006 | | -0.005 | | | 0.104 | | 0.123 |
| | | | (0.006) | | (0.004) | | | (0.090) | | (0.085) |
| IIT | | | | 1.918*** | 1.837*** | | | | -0.063 | 0.005 |
| | | | | (0.511) | (0.516) | | | | (0.133) | (0.139) |
| $M^G$.IIT | | | | -0.005# | -0.004 | | | | 0.010** | -0.004 |
| | | | | (0.003) | (0.003) | | | | (0.003) | (0.010) |
| Constant | 22.647*** | 22.644*** | 22.731*** | 21.801*** | 21.882*** | 21.655*** | 21.608*** | 21.738*** | 21.655*** | 21.728*** |
| | (0.470) | (0.491) | (0.471) | (0.488) | (0.480) | (0.160) | (0.150) | (0.155) | (0.202) | (0.221) |
| # Observations | 120 | 120 | 120 | 120 | 120 | 138 | 138 | 99 | 138 | 99 |
| df_m | 23 | 24 | 26 | 26 | 28 | 26 | 27 | 26 | 29 | 28 |
| r2 | 0.965 | 0.965 | 0.967 | 0.975 | 0.975 | 0.996 | 0.996 | 0.998 | 0.996 | 0.998 |
| Year fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

**Note:** (1) Legend: # p<.1; * p<.05; ** p<.01; *** p<.001 (2) Standard errors reported in brackets.