Protection for Sale Under Monopolistic Competition: An Empirical Investigation

Pao-Li Chang*

Myoung-Jae Lee

School of Economics and Social Sciences

Singapore Management University, Singapore 259756

June 14, 2005

Abstract

Previous cross-sectional empirical works based on Grossman and Helpman (1994) have often adopted homogeneous (perfect competition) market structure. Recently, Chang (2005) suggested that the endogenous protection structure under the political framework of Grossman and Helpman (1994) changes systematically when the underlying market structure is monopolistically competitive. In this paper, we adopt a general empirical specification that accommodates both monopolistically and perfectly competitive sectors. Our results favor the general specification of heterogeneous market structures over the homogeneous perfect competition specification. This empirical finding thus suggests an intricate relationship between import protection and import penetration: it depends on whether the sector is perfectly competitive or monopolistically competitive and whether the sector is politically organized or not.

Keywords: endogenous trade policy; campaign contribution; monopolistic competition; intraindustry trade; import penetration

JEL classification: F12; F13

*Tel: +65-6828-0830; fax: +65-6828-0833. *E-mail address:* plchang@smu.edu.sg (P.-L. Chang).

1. INTRODUCTION

Previous cross-sectional empirical works based on Grossman and Helpman (1994) (henceforth G-H) have often adopted homogeneous (perfect competition) market structure. Recently, Chang (2005) suggests that the endogenous protection structure under the political framework of G-H changes systematically when the underlying market structure is characterized by monopolistic competition. First, the benchmark welfare-maximizing import tariff is strictly positive, in contrast with the free trade prediction under perfect competition. Secondly, lobbying efforts spent by competing interest groups, each representing the owners of a factor specific to a sector, then raise import protection levels in organized sectors (represented by a lobby) and lower them in unorganized sectors (not represented by a lobby), relative to the benchmark levels. The endogenous tariff level, nonetheless, will not fall below zero in unorganized sectors. This is contrary to G-H where unorganized sectors will be penalized by negative protection. Lastly, the level of import protection is predicted in Chang (2005) to vary inversely with the degree of import penetration, regardless of whether the sector is organized or not. In other words, sectors with higher import penetration will receive lower protection. This negative relationship applies only to organized sectors in G-H; in unorganized sectors, higher import penetration leads to higher protection (or more precisely, less negative protection).

The last theoretical prediction is especially of interest to empirical researchers. The general perception before the G-H model was that a sector badly injured by imports would tend to receive higher protection. This perception was supported by most empirical works conducted before G-H; however, their findings were often derived using ad hoc regression models without theoretical underpinnings. The G-H model presented a challenge to this conventional view, as it predicts the opposite for organized sectors. Several authors (Goldberg and Maggi, 1999; Gawande and Bandyopadhyay, 2000; Eicher and Osang, 2002) have hence estimated the G-H model and put this prediction to test. In general, their findings are supportive of the G-H model. The results of Chang (2005) offer a picture completely at odd with the conventional view, as the import protection actually falls with import penetration, regardless of whether the sector is organized or not. This difference in findings between G-H and Chang (2005) suggests that the predictions of the G-H-type models in this regard will depend on the nature of the market structure. Hence, a correctly-specified empirical model relating to the G-H model should allow for heterogeneous responses of protection to import penetration across sectors depending on the market structure. This has not been considered in previous cross-sectional G-H-type empirical works.

Our paper attempts to account for this heterogeneity empirically. To do this, we adopt the G-H model for perfectly competitive sectors and the Chang (2005) model for monopolistically competitive sectors. We examine whether the Chang (2005) model is borne out by the data, and whether the general specification allowing both market structures provides a better fit to the data than the homogeneous perfect competition model. If the answers are positive, we may thus accept the estimates of the underlying political parameters derived from the general specification as being more representative of the economy under study than suggested by previous estimates derived from homogeneous perfect competition models.

The rest of the paper is organized as follows. In Section 2, we review the basic elements of G-H and Chang (2005) and integrate the two models. In Section 3, we present the empirical model for the general specification allowing heterogeneous market structures and explain the estimation methodology. The general specification is then compared against the homogeneous perfect competition specification based on non-nested hypothesis tests. The empirical results are given in Section 3.5. Finally, Section 4 concludes. Technical notes regarding the estimation methodologies are given in the appendices.

2. THE PROTECTION-FOR-SALE MODEL WITH HETEROGENEOUS MARKET STRUCTURE

In this section, we integrate the original Protection-for-Sale model by Grossman and Helpman (1994) and the modified model by Chang (2005). This construct allows the market structure to be potentially heterogeneous across sectors, either perfectly or monopolistically competitive. Suppose that a country is populated by individuals with identical preferences, given by:

$$U = C_0 + \sum_{i=1}^{n} U_i(C_i) \tag{1}$$

where C_0 denotes consumption of good 0, C_i denotes consumption of good i, $i = 1, 2, \ldots, n$, and U_i is an increasing concave function. Good 0 serves as numeraire, with a world and domestic price equal to 1. Let P_i denote the domestic price of good i. The demand for good i implied by the preferences in (1) is denoted $D_i(P_i)$, where $D_i(\cdot)$ is the inverse of $U_i'(\cdot)$. The indirect utility of an individual with income E is given by $V = E + \sum_{i=1}^{n} S_i(P_i)$, where $S_i(P_i) = U_i(D_i(P_i)) - P_i D_i(P_i)$ is the consumer surplus derived from good i. If sector i is monopolistically competitive, C_i represents the aggregate consumption of differentiated goods in sector i, with the aggregation following the Dixit-Stiglitz functional form (Dixit and Stiglitz, 1977):

$$C_i = (\sum_{k=1}^{m_i} c_{hi,k}^{\varrho_i} + \sum_{k=1}^{m_i^*} c_{fi,k}^{\varrho_i})^{\frac{1}{\varrho_i}} \qquad 0 < \varrho_i < 1 \tag{2}$$

where $c_{hi,k}$ ($c_{fi,k}$) is the consumption of home (foreign) variety k of good i and m_i (m_i^*) is the number of varieties of good i produced at home (abroad). The corresponding aggregate price level P_i for the differentiated goods in sector i is:

$$P_i = (\sum_{k=1}^{m_i} p_{hi,k}^{1-\sigma_i} + \sum_{k=1}^{m_i^*} p_{fi,k}^{1-\sigma_i})^{\frac{1}{1-\sigma_i}} \tag{3}$$

where $p_{hi,k}$ ($p_{fi,k}$) is the consumer price for home (foreign) variety k of good i, and $\sigma_i = \frac{1}{1-\varrho_i} > 1$ is the elasticity of substitution among different varieties of good i.

Good 0 is taken to be a homogeneous good, produced one-to-one from labor, and traded freely and costlessly, so that the wage is equal to one at home and abroad. Production of the other goods requires labor and a sector-specific input. The various specific inputs are available in inelastic supply $\bar{K}_i$, $i = 1, 2, \ldots, n$. If sector i is perfectly competitive, the production technology exhibits constant returns to scale. If sector i is monopolistically competitive, each variety of the differentiated goods is assumed to require a fixed amount of the sector-specific factor k_i in order to produce at all; after that, there is a constant unit labor requirement θ_i. Thus, the number of varieties produced at home in sector i is $m_i = \bar{K}_i / k_i$. The technology abroad to produce the differentiated products is assumed to be the same as at home. Thus, the number of varieties produced abroad in the same sector is $m_i^* = \bar{K}_i^* / k_i$, where $\bar{K}_i^*$ is the amount of sector-specific factor i the foreign country is endowed with.

Let the domestic import policy τ_i denote one plus the ad valorem import tariff rate and the domestic export policy s_i represent one plus the ad valorem export subsidy rate for sector i. Similarly, let τ_i^* and s_i^* represent the corresponding foreign import and export policy for sector i, defined in a similar way.

Suppose sector i is perfectly competitive and the exogenous world price is P_i^*. Then, the domestic price is $P_i = \tau_i P_i^*$ in an import-competing sector, and is $P_i = s_i P_i^*$ in an export sector. The returns to specific factor i depend only on P_i and are denoted by $\Pi_i(P_i)$. It follows that the supply function of good i is $Y_i(P_i) = \Pi_i^{'}(P_i)$.

Suppose sector i is monopolistically competitive instead. Assume that there are a large number of varieties (home and foreign combined) available to the consumer in sector i. Given the preferences specified in (2), each variety's producer faces an approximately constant elasticity of demand, σ_i. Thus, with profit maximization, the producer of each variety charges the same price: $p_{hi,k} = p_{hi} = p_{fi}^* = \frac{\theta_i \sigma_i}{\sigma_i - 1}$, where p_{fi}^* is the producer (and consumer) price for each foreign variety of good i in the foreign market. Since the producer prices of the home and the foreign variety are the same, the difference in the consumer prices of the home and the foreign variety in the domestic market reflects the government interventions in trade: $p_{fi} = \frac{\tau_i}{s_i^*} p_{hi}$. Thus, the aggregate price index for differentiated good i in (3) can be simplified as:

$$P_i = p_{hi}(m_i + m_i^*(\frac{\tau_i}{s_i^*})^{1-\sigma_i})^{\frac{1}{1-\sigma_i}}. \quad (4)$$

The utility function U_i in (1) for monopolistically competitive sectors is assumed to take the functional form: $U_i = E_i \, lnC_i$. This amounts to assuming that an individual allocates a fixed amount of expenditure E_i on good i. The rest of the world is assumed to share the same preference structure, but with a possibly different allocation of expenditure on various goods, E_i^*. Given this, we can derive the demand for a representative home and foreign variety of good i as:

$$c_{hi} = \frac{E_i}{p_{hi}} \frac{1}{m_i + m_i^*(\frac{\tau_i}{s_i^*})^{1-\sigma_i}} \quad (5)$$
$$c_{fi} = \frac{E_i}{p_{fi}} \frac{(\frac{\tau_i}{s_i^*})^{1-\sigma_i}}{m_i + m_i^*(\frac{\tau_i}{s_i^*})^{1-\sigma_i}}.$$

Similarly, a foreign individual will consume a representative home and foreign variety of good i

according to:

$$
\begin{aligned}
c^*_{hi} &= \frac{E^*_i}{p^*_{hi}} \; \frac{(\frac{\tau^*_i}{s_i})^{1-\sigma_i}}{m_i(\frac{\tau^*_i}{s_i})^{1-\sigma_i} + m^*_i} \\
c^*_{fi} &= \frac{E^*_i}{p^*_{fi}} \; \frac{1}{m_i(\frac{\tau^*_i}{s_i})^{1-\sigma_i} + m^*_i}
\end{aligned}
\tag{6}
$$

where $p^*_{hi} = \frac{\tau^*_i}{s_i} p^*_{fi}$ is the consumer price abroad for a representative home variety of good i. Given the domestic and foreign demand for its product, a representative home producer of differentiated good i will produce at the scale of $(Nc_{hi} + N^* c^*_{hi})$, where N (N^*) is the total home (foreign) population. Thus, the returns to the specific factor used in differentiated sector i are: $\Pi_i(\tau_i, s_i) = m_i(p_{hi} - \theta_i)(Nc_{hi} + N^* c^*_{hi})$.

The net tariff revenue from sector i, expressed on a per capita basis, is given by:

$$
\begin{aligned}
R_i =& (1 - I^m_i)(P_i - P^*_i)[C_i - \frac{1}{N} X_i] \\
&+ I^m_i [m^*_i(\tau_i - 1)\frac{p_{hi}}{s^*_i} \; c_{fi} - \frac{N^*}{N} m_i (1 - \frac{1}{s_i}) p_{hi} \; c^*_{hi}]
\end{aligned}
\tag{7}
$$

where $X_i = Y_i(P_i)$ is the domestic aggregate output of good i in a perfectly competitive sector, and I^m_i is an indicator variable, which equals one if sector i is monopolistically competitive and zero otherwise. It is assumed that the government redistributes the revenue $R = \sum_{i=1}^n R_i$ evenly to each individual.

Summing indirect utilities over all individuals, and noting that aggregate income is the sum of labor income, returns to specific factors and tariff revenue, one obtains aggregate welfare:

$$
W = N + \sum_{i=1}^n \Pi_i + N \sum_{i=1}^n (R_i + S_i). \tag{8}
$$

We now describe the political structure. Suppose that in some subset of sectors $L \subset \{1, 2, \cdots n\}$, the specific-factor owners are able to form a lobby. Let α_i denote the fraction of population that owns specific factor i. Assume that each individual owns a unit of labor and at most one type of specific factor. Summing indirect utilities over all individuals who belong to lobby i, we obtain lobby i's aggregate well-being:

$$W_i = \alpha_i N + \Pi_i + \alpha_i N \sum_{i=1}^{n} (R_i + S_i). \tag{9}$$

Lobbies compete noncooperatively for the government's favor and propose contribution schedules, $\mathbb{C}_i(\tau, s)$, contingent on the trade-policy vector set by the government, (τ, s). Lobby i's objective is to maximize the net aggregate well-being given by $W_i - \mathbb{C}_i$. Given the contribution schedules offered by the lobbies, the government in turn selects a trade-policy vector (τ, s) to maximize its politically-motivated objective function, which is a combination of welfare and contributions:

$$G = \sum_{i \in L} \mathbb{C}_i + aW \tag{10}$$

where $a \geq 0$ captures the weight of welfare in the government's objective relative to campaign contributions. As shown in G-H, if the contribution schedules offered by lobbies are truthful, the government's objective function in (10) is equivalent to:

$$\tilde{G} = \sum_{i \in L} W_i + aW. \tag{11}$$

To find the equilibrium trade policy, one can rewrite $\tilde{G}$ as:

$$\tilde{G} = (a + \alpha_L)N + \sum_{i=1}^{n} (a + I_i)\Pi_i + (a + \alpha_L)N \sum_{i=1}^{n} (R_i + S_i). \tag{12}$$

where $\alpha_L \equiv \sum_{i \in L} \alpha_i$ denotes the fraction of population that is represented by a lobby, and I_i is an indicator variable that equals one if sector i is organized and zero otherwise. We focus on the endogenous import policy henceforth, and refer interested readers to Chang (2005) for details on the endogenous export policy. Let $t_i \equiv \tau_i - 1$ denote the ad valorem import tariff rate. Taking the first-order derivative with respect to (12) yields the following results:

PROPOSITION 1 (Endogenous Protection Structure) *If the contribution schedules of the lobbies are truthful, the import policy that will emerge in the political equilibrium for a perfectly competitive sector satisfies*

$$\frac{t_i^o}{1+t_i^o} = \frac{I_i - \alpha_L}{a + \alpha_L}\ \frac{z_i^o}{e_i^o} \tag{13}$$

where $z_i^o = X_i^o / M_i^o$ *is the equilibrium ratio of domestic output to imports, and* $e_i^o = -M_i^{o\prime} P_i^o / M_i^o$ *is the elasticity of import demand.*

On the other hand, for a monopolistically competitive sector, the import policy that will emerge in the political equilibrium satisfies

$$\frac{t_j^o}{1+t_j^o} = \frac{I_j + a}{a + \alpha_L}\ \frac{\frac{\sigma_j - 1}{\sigma_j}}{\sigma_j + \frac{1}{\tilde{z}_j^o}} \tag{14}$$

where $\tilde{z}_j^o = p_{hj} m_j c_{hj}^o \,/\, p_{fj}^o m_j^* c_{fj}^o$ *is the equilibrium ratio of domestic output supplied to the domestic market relative to imports, and* σ_j *is the constant elasticity of substitution between domestic output and imports, or approximately, the elasticity of import demand.*

Proposition 1 indicates that for a perfectly competitive sector, import protection decreases with higher import penetration ($1/z_i$) if the sector is organized, and increases with higher import penetration if the sector is unorganized. On the other hand, for a monopolistically competitive sector, import protection always decreases with higher import penetration ($1/\tilde{z}_i$), regardless of whether the sector is organized or not. Next, Proposition 1 indicates that the effects of the political parameters, α_L and a, are comparable under perfectly and monopolistically competitive sectors. As more people are politically represented (a larger α_L), the general protection level decreases under both perfectly and monopolistically competitive sectors. Moreover, as the government becomes more concerned with aggregate welfare (a larger a), the endogenous protection level approaches the welfare-maximizing level, which is free trade for a perfectly competitive sector and a positive import tariff $\frac{(\sigma_j - 1)/\sigma_j}{\sigma_j + 1/\tilde{z}_j^o}$ for a monopolistically competitive sector.

3. THE ECONOMETRIC MODEL

If heterogeneity in market structure is ignored and all sectors are taken to be perfectly competitive, equation (13) as derived originally by Grossman and Helpman (1994) applies to all sectors. This theoretical specification has been the basis of previous empirical researches such as Goldberg and Maggi (1999), Gawande and Bandyopadhyay (2000), and Eicher and Osang (2002). For example, Goldberg and Maggi (1999) form the following structural equation from (13):

$$\begin{aligned} \frac{t_i}{1+t_i} e_i &= \frac{I_i - \alpha_L}{a + \alpha_L} z_i + \epsilon_{p,i} \\ &= \gamma_p z_i + \delta_p I_i z_i + \epsilon_{p,i} \end{aligned} \tag{15}$$

where $\gamma_p = -\alpha_L/(a+\alpha_L)$ and $\delta_p = 1/(a+\alpha_L)$. We will call this specification the homogeneous perfect competition (PC) model. Theory (13) implies that the parameters under the PC model should have the following signs: (i) $\gamma_p < 0$, (ii) $\delta_p > 0$, (iii) $\gamma_p + \delta_p > 0$.

On the other hand, if all sectors are taken to be monopolistically competitive, equation (14) applies to all sectors. A structural equation based on (14) can be derived in parallel with (15) as:

$$\begin{aligned} \frac{t_i}{1+t_i} e_i &= \frac{I_i + a}{a + \alpha_L} \frac{e_i - 1}{e_i + \frac{1}{\tilde{z}_i}} + \epsilon_{m,i} \\ &= \gamma_m w_i + \delta_m I_i w_i + \epsilon_{m,i} \end{aligned} \tag{16}$$

where $\gamma_m = a/(a+\alpha_L)$, $\delta_m = 1/(a+\alpha_L)$, and $w_i = \frac{e_i - 1}{e_i + \frac{1}{\tilde{z}_i}}$. Note that we have used e_i as a proxy for σ_i in (16), as σ_i being the elasticity of substitution among different varieties of good i is also the elasticity of substitution between home and foreign varieties, which is approximately the elasticity of import demand, e_i. Theory (14) implies that the parameters for monopolistically competitive sectors should have the following signs: (i) $\gamma_m > 0$, (ii) $\delta_m > 0$, (iii) $\gamma_m + \delta_m > 1$.

In this paper, we consider a generalized specification of endogenous protection structure based on Proposition 1, which accommodates both potential market structures:

$$\frac{t_i}{1+t_i} e_i = (1 - I_i^m)\{\gamma_p z_i + \delta_p I_i z_i + \epsilon_{p,i}\} + I_i^m \{\gamma_m w_i + \delta_m I_i w_i + \epsilon_{m,i}\}, \tag{17}$$

subject to the constraints $-\gamma_p + \gamma_m = 1$ and $\delta_p = \delta_m$. We will call this generalized specification the heterogeneous market structure (HC) model. Altogether, the parameters of interest under the HC model should observe the following signs: (i) $\gamma_p < 0$, (ii) $\delta_p > 0$, (iii) $\gamma_p + \delta_p > 0$, (iv) $\gamma_m > 0$, (v) $\delta_m > 0$, (vi) $\gamma_m + \delta_m > 1$. We are interested to learn whether the HC model provides a statistically better fit to the data than the PC model frequently used in previous empirical studies, and if this is the case, what the implied values are for the underlying political parameters a and α_L given the estimated HC model.

3.1 Data and Measurement

We briefly discuss the measurement and endogeneity issues associated with estimation of the specifications (15) and (17). The data set we used was kindly provided by Eicher and Osang (2002), which is a reconstruction of the data set described in Goldberg and Maggi (1999). We refer the reader to these two sources for detailed explanations of the data set. In brief, the sectors investigated are 106 United States manufacturing sectors at the 3-digit SIC level. The coverage ratios for nontariff barriers (NTB) in year 1983, instead of tariffs, are used to proxy for the noncooperative endogenous protection level predicted by the above theories. This is in view of the fact that tariff levels in reality are set cooperatively among nations in international trade negotiations. Since the coverage ratio can only take values between 0 and 1, the protection variable t_i is censored. We follow the benchmark mapping in Goldberg and Maggi (1999) between the latent protection level and the nontariff barrier, that is, for coverage ratios less than 1, they reflect the equivalent tariff level. For latent equivalent tariffs higher than 100 percent, they are mapped to the coverage ratio 1.

The trade elasticity estimates from Shiells et al. (1986) are used to measure both the import elasticity in (15) for perfectly competitive sectors, and the elasticity of substitution supposedly in (16) for monopolistically competitive sectors. The import elasticity measure was brought to the left-hand side of the estimating equation for perfectly competitive sectors in Goldberg and Maggi (1999), because the variable is endogenous by theory (13) and yet no suitable instrumental variables could be clearly identified. The elasticities of substitution for monopolistically competitive sectors, however, are not endogenous by theory (14). To facilitate estimation and comparison, however, we also phrase the structural equation for monopolistically competitive sectors in a similar fashion such that the left-hand side variable is also a composite of the protection measure and the elasticity measure as in perfectly competitive sectors. The elasticity terms that remain on the right-hand side of the structural equation (16) for monopolistically competitive sectors can be taken as exogenous by theory (14).

The inverse import-penetration ratio z_i for perfectly competitive sectors in (15) is measured by the ratio of the value of shipments over imports as in Goldberg and Maggi (1999). The alternative import-penetration ratio $\tilde{z}_i$ for monopolistically competitive sectors in (16), however, requires slight modifications. It is measured as the ratio of the value of shipment to the domestic market over imports (= (value of shipment - exports)/imports). Because both the import and domestic production level depend on the protection level, these two inverse penetration ratios are endogenous. We follow Goldberg and Maggi (1999) in the selection of the explanatory variables for the inverse penetration ratio, z_i. These include physical capital, inventories, engineers/scientists, white-collar worker, skilled labor, semiskilled labor, cropland, pasture, forest, coal, petroleum, and minerals (which are the explanatory variables used for import penetration in Trefler, 1993), as well as seller concentration, buyer concentration, seller number of firms, buyer number of firms, scale, capital stock, unionization, geographic concentration, and tenure (which are a subset of the explanatory variables used for import protection in the same source). This collection of explanatory variables are also used for the composite inverse penetration ratio, w_i.

The political-organization dummy I_i is exogenous in theory, but is empirically measured based on sectoral political contributions. Goldberg and Maggi (1999) construct the dummy using the threshold level of \$1,000,000,000 for political action committee (PAC) contributions. A sector is considered politically organized for trade-policy purposes, if the sector's PAC contribution is above the threshold. By constructing the organization dummy in such a way implies that I_i is potentially endogenous since the political contribution level $\mathbb{C}_i$ is endogenous. However, we will treat I_i as exogenous, because allowing I_i to be endogenous poses significant methodological problems, the technical details of which are explained in Appendix A.

To classify sectors as perfectly or monopolistically competitive, we begin with a simplistic criterion suggested by the data based on the elasticity measure. For theory (14) to hold for monopolistically competitive sectors, the elasticity of substitution must be greater than one. Thus, sectors with trade elasticity measures less than or equal to one are classified as perfectly competitive ($I_i^m = 0$ if $e_i \leq 1$), while sectors with elasticity measures larger than one are classified as monopolistically competitive ($I_i^m = 1$ if $e_i > 1$). We then explore alternative measures of I_i^m based on a further criterion that for a sector to be classified as monopolistically competitive, the degree of seller competition in the sector should be low enough (such that individual firms maintain some degree of monopoly power). We use one of the exogenous variables, the seller number of firms, to measure the degree of seller competition. Specifically, the seller number of firms is the number of companies in a sector scaled by the sector's total sales. Let '*scomp*' denote the degree of seller competition and $\overline{\kappa}$ the chosen threshold level. Then, the refined market structure dummy is defined as follows: a sector is considered monopolistically competitive, if the elasticity measure is greater than one and if the sector faces relatively low degrees of competition (that is, $I^m = 1$, if $e_i > 1$ and $scomp_i \leq \overline{\kappa}$). We consider an extensive range of $\overline{\kappa}$, starting with the highest observed value of '*scomp*', which is equivalent to imposing no extra restriction, to the lowest value possible that still gives rise to a tolerable degree of freedom for the monopolistically competitive sectors.

3.2 The Full Econometric Model

Let SF stand for 'structural form'. Given the above discussion, the full econometric model for the PC model we consider is, with $\bar{c}_i = \frac{1}{2} e_i > 0$,

$$\begin{aligned}
y \text{ SF} \quad &: \quad y_i = \max\{0, \min(y_i^*, \bar{c}_i)\} \\
y^* \text{ SF} \quad &: \quad y_i^* = \frac{t_i^*}{1+t_i^*} e_i = \gamma_p z_i + \delta_p I_i z_i + \epsilon_{p,i} \\
z_i \quad &= \quad \zeta_p' Z_i + u_{p,i} \\
\text{observed} \quad &: \quad y_i,\ z_i,\ I_i,\ Z_i,\ e_i,\ \bar{c}_i,\ i = 1, ..., n,\ iid \text{ across } i.
\end{aligned}$$

where $u_{p,i}$ and $\epsilon_{p,i}$ are dependent, which makes z_i endogenous in y SF. The latent variable t_i^* indicates the true level of protection, which is censored at zero and one when measured by the NTB coverage ratio. It follows that the upper censoring point for y^* SF is $\bar{c}_i$, which varies across i. The vector Z_i consists of one and the explanatory variables for the inverse penetration ratio z_i. The HC model, on the other hand, takes the following form:

$$\begin{aligned}
y \text{ SF} \quad &: \quad y_i = \max\{0, \min(y_i^*, \bar{c}_i)\} \\
y^* \text{ SF} \quad &: \quad y_i^* = \frac{t_i^*}{1+t_i^*} e_i = (1 - I_i^m)\{\gamma_p z_i + \delta_p I_i z_i + \epsilon_{p,i}\} + I_i^m\{\gamma_m w_i + \delta_m I_i w_i + \epsilon_{m,i}\} \\
z_i \quad &= \quad (1 - I_i^m)\{\zeta_{pp}' Z_i + u_{pp,i}\} + I_i^m\{\zeta_{pm}' Z_i + u_{pm,i}\} \\
w_i \quad &= \quad (1 - I_i^m)\{\zeta_{mp}' Z_i + u_{mp,i}\} + I_i^m\{\zeta_{mm}' Z_i + u_{mm,i}\} \\
\text{observed} \quad &: \quad y_i,\ z_i,\ w_i,\ I_i,\ I_i^m,\ Z_i,\ e_i,\ \bar{c}_i,\ i = 1, ..., n,\ iid \text{ across } i.
\end{aligned}$$

where $u_{pp,i}$ and $\epsilon_{p,i}$ are dependent, and so are $u_{mm,i}$ and $\epsilon_{m,i}$, which makes z_i and w_i endogenous. In this model, heterogeneity in market structure across sectors is explicitly taken into account. In particular, we allow the protection structure y_i^*, the inverse penetration ratio z_i, and the composite inverse penetration ratio w_i to behave differently in perfectly competitive sectors ($I_i^m = 0$) from in monopolistically competitive sectors ($I_i^m = 1$), with y^* SF explicitly guided by Proposition 1. Without any further constraints, this system of specifications would be equivalent to two independent systems of specifications, one for each group of sectors belonging to the same market structure. However, as discussed earlier, the SF parameters under the two market structures are related in the HC model, through the constraints $-\gamma_p + \gamma_m = 1$ and $\delta_p = \delta_m$.

3.3 Estimation

In view of the *iid* assumption, the subscript i will be omitted often in the following. The goal is to estimate the SF parameters γ_p and δ_p in the PC model, and γ_p, δ_p, γ_m, and δ_m in the HC model subject to the constraints $-\gamma_p + \gamma_m = 1$ and $\delta_p = \delta_m$. There are two econometric problems in pursuing the goal. One is the censoring in the y SF and the other is the endogeneity of z and w. The z- (and w-) equation regressor vector Z that includes unity is essentially an instrument vector for z (and w), but due to the nonlinearity caused by the censoring, the usual instrumental variable estimator (IVE) for linear models is not applicable. As explained in Appendix A, we explored four methods. They differ in terms of whether the censored maximum likelihood estimator (CMLE) or the censored least absolute deviation estimator (CLAD) is used to estimate the y SF, and whether the minimum distance estimator (MDE) or the two-stage least square type estimator (2SLS) is used to account for the endogeneity of the z- (and w-) equation. The system maximum likelihood estimator might be yet another option. Given the relatively small sample size coupled with the relatively large number of parameters, we are led to adopt the 2SLS-CMLE combination after a preliminary data analysis, as it is less restrictive than the system MLE in terms of the requisite assumptions, and is more numerically stable than the MDE/CLAD.

Let σ_a and $\rho_{a,b}$ stand for, respectively, the standard error of variable a and the correlation between variables a and b. For the PC model, substituting the z- equation into the y^* SF, we get the y^* SF-2SLS for the PC model and the conditional variance of the error term:

$$y^* \text{ SF-2SLS} \quad : \quad y^* = \gamma_p \zeta_p' Z + \delta_p I \zeta_p' Z + \{(\gamma_p + \delta_p I) u_p + \epsilon_p\}, \tag{18}$$

$$\sigma_{v_p}^2 \quad \equiv \quad V(\{\cdot\}|Z, I) = (\gamma_p + \delta_p I)^2 \sigma_{u_p}^2 + 2(\gamma_p + \delta_p I)\rho_{u_p,\epsilon_p}\sigma_{u_p}\sigma_{\epsilon_p} + \sigma_{\epsilon_p}^2. \tag{19}$$

The error term in $\{\cdot\}$ is heteroscedastic depending on Z and I, and consists of two errors. Let $\hat{\zeta}_p$ and $\hat{\sigma}_{u_p}$ denote the first-stage least square estimate (LSE) for ζ_p and σ_{u_p} in the z- equation. In the second stage, the y SF-2SLS can be estimated applying the usual CMLE with ζ_p and σ_{u_p} replaced by $\hat{\zeta}_p$ and $\hat{\sigma}_{u_p}$, and with an adjustment owing to the heteroscedasticity. The second-stage log-likelihood function for the y SF-2SLS of the PC model is

$$\begin{aligned} L_p = \sum_i \{ \ 1[y_i &= 0] \ln \Phi(\frac{-\gamma_p \hat{\zeta}_p' Z_i - \delta_p I_i \hat{\zeta}_p' Z_i}{\sigma_{v_{p,i}}}) \\ +1[0 &< y_i < \bar{c}_i] \ln \frac{\phi((y_i - \gamma_p \hat{\zeta}_p' Z_i - \delta_p I_i \hat{\zeta}_p' Z_i)/\sigma_{v_{p,i}})}{\sigma_{v_{p,i}}} \\ +1[y_i &= \bar{c}_i] \ln \Phi(\frac{\gamma_p \hat{\zeta}_p' Z_i + \delta_p I_i \hat{\zeta}_p' Z_i - \bar{c}_i}{\sigma_{v_{p,i}}}) \ \}, \end{aligned} \tag{20}$$

which is to be maximized for γ_p, δ_p, σ_{ϵ_p}, and ρ_{u_p,ϵ_p}; note that $\sigma_{v_{p,i}}$ is a function of Z_i, I_i and the parameters, which are suppressed to simplify notations. Denote the resulting estimates $(\dot{\gamma}_p, \dot{\delta}_p, \dot{\sigma}_{\epsilon_p}, \dot{\rho}_{u_p,\epsilon_p})$.

Analogous steps are needed for the HC model. Substituting the z- and w- equation into the y^* SF, we get the y^* SF-2SLS for the HC model and the conditional variance of the error term:

$$\begin{aligned} y^* \text{ SF-2SLS} \quad &: \quad y^* = (1 - I^m)(\gamma_p \zeta_{pp}' Z + \delta_p I \zeta_{pp}' Z) + I^m(\gamma_m \zeta_{mm}' Z + \delta_m I \zeta_{mm}' Z) \\ &\quad +\{(1 - I^m)(\gamma_p + \delta_p I) u_{pp} + I^m(\gamma_m + \delta_m I) u_{mm} + (1 - I^m)\epsilon_p + I^m \epsilon_m\}, \end{aligned} \tag{21}$$

$$\begin{aligned} \sigma_{v_h}^2 \quad &\equiv \quad V(\{\cdot\}|Z, I, I^m) = (1 - I^m)[(\gamma_p + \delta_p I)^2 \sigma_{u_{pp}}^2 + 2(\gamma_p + \delta_p I)\rho_{u_{pp},\epsilon_p}\sigma_{u_{pp}}\sigma_{\epsilon_p} + \sigma_{\epsilon_p}^2] \\ &\quad +I^m[(\gamma_m + \delta_m I)^2 \sigma_{u_{mm}}^2 + 2(\gamma_m + \delta_m I)\rho_{u_{mm},\epsilon_m}\sigma_{u_{mm}}\sigma_{\epsilon_m} + \sigma_{\epsilon_m}^2]. \end{aligned} \tag{22}$$

The error term in $\{\cdot\}$ is heteroscedastic depending on Z, I, I^m, and consists of four errors. Denoting the first-stage LSE for ζ_{pp}, $\sigma_{u_{pp}}$, ζ_{mm}, and $\sigma_{u_{mm}}$ in the z- and w- equations as $\hat{\zeta}_{pp}$, $\hat{\sigma}_{u_{pp}}$, $\hat{\zeta}_{mm}$, and

$\hat{\sigma}_{u_{mm}}$, the resulting second-stage log-likelihood function for the y SF-2SLS of the HC model is

$$
\begin{aligned}
L_h = \sum_i \{ \; 1[y_i &= 0] \ln \Phi\Big(\frac{-(1-I^m)(\gamma_p \hat{\zeta}_{pp}' Z + \delta_p I \hat{\zeta}_{pp}' Z) - I^m(\gamma_m \hat{\zeta}_{mm}' Z + \delta_m I \hat{\zeta}_{mm}' Z)}{\sigma_{v_{h,i}}}\Big) \\
+1[0 &< y_i < \bar{c}_i] \ln \frac{\phi((y_i - (1-I^m)(\gamma_p \hat{\zeta}_{pp}' Z + \delta_p I \hat{\zeta}_{pp}' Z) - I^m(\gamma_m \hat{\zeta}_{mm}' Z + \delta_m I \hat{\zeta}_{mm}' Z))/\sigma_{v_{h,i}})}{\sigma_{v_{h,i}}} \\
+1[y_i &= \bar{c}_i] \ln \Phi\Big(\frac{(1-I^m)(\gamma_p \hat{\zeta}_{pp}' Z + \delta_p I \hat{\zeta}_{pp}' Z) + I^m(\gamma_m \hat{\zeta}_{mm}' Z + \delta_m I \hat{\zeta}_{mm}' Z) - \bar{c}_i}{\sigma_{v_{h,i}}}\Big) \; \},
\end{aligned}
\tag{23}
$$

which is to be maximized for γ_p, δ_p, σ_{ϵ_p}, γ_m, δ_m, σ_{ϵ_m}, ρ_{u_{pp},ϵ_p}, and ρ_{u_{mm},ϵ_m}, subject to the constraints $-\gamma_p + \gamma_m = 1$ and $\delta_p = \delta_m$. Denote the resulting estimates $(\ddot{\gamma}_p, \ddot{\delta}_p, \ddot{\sigma}_{\epsilon_p}, \ddot{\gamma}_m, \ddot{\delta}_m, \ddot{\sigma}_{\epsilon_m}, \ddot{\rho}_{u_{pp},\epsilon_p}, \ddot{\rho}_{u_{mm},\epsilon_m})$. The details of how to account for the effect of first-stage estimation errors on the second-stage asymptotic variance are explained in Appendix B.

3.4 Hypothesis Testing

To test the validity of the two competing models, we apply the *J-test* in Davidson and MacKinnon (1981) for non-nested hypotheses. Define $f(Z, I, \gamma_p, \delta_p, \zeta_p) \equiv \gamma_p \zeta_p' Z + \delta_p I \zeta_p' Z$, which is the SF-2SLS regression function of the PC model in (18). Similarly, define $g(Z, I, I^m, \gamma_p, \delta_p, \gamma_m, \delta_m, \zeta_{pp}, \zeta_{mm}) \equiv (1-I^m)(\gamma_p \zeta_{pp}' Z + \delta_p I \zeta_{pp}' Z) + I^m(\gamma_m \zeta_{mm}' Z + \delta_m I \zeta_{mm}' Z)$, which is the SF-2SLS regression function of the HC model in (21). Then, the J-test for the PC model as the null hypothesis proceeds as follows. First, create an artificially augmented regression function of f:

$$\tilde{f} \equiv f(Z, I, \gamma_p, \delta_p, \zeta_p) + \mu_h \hat{g},$$

where $\hat{g} = g(Z, I, I^m, \ddot{\gamma}_p, \ddot{\delta}_p, \ddot{\gamma}_m, \ddot{\delta}_m, \hat{\zeta}_{pp}, \hat{\zeta}_{mm})$ is the fitted value based on the parameter estimates from the HC model. Next, obtain the 2SLS-CMLE based on a modified likelihood function $\tilde{L}_p$ as in (20) but with the augmented regression function $\tilde{f}$. Tests for $\mu_h = 0$. If $\mu_h = 0$ is rejected, then the PC model is rejected to the direction of the HC model.

The J-test for the HC model as the null hypothesis can be done analogously. Form an artificially augmented regression function of g:

$$\tilde{g} \equiv g(Z, I, I^m, \gamma_p, \delta_p, \gamma_m, \delta_m, \zeta_{pp}, \zeta_{mm}) + \mu_p \hat{f},$$

where $\hat{f} = f(Z, I, \dot{\gamma}_p, \dot{\delta}_p, \hat{\zeta}_p)$ is the fitted value based on the parameter estimates from the PC model. Obtain the 2SLS-CMLE based on a modified likelihood function $\tilde{L}_h$ as in (23) but with the augmented regression function $\tilde{g}$. Tests for $\mu_p = 0$. If $\mu_p = 0$ is rejected, then the HC model is rejected to the direction of the PC model.

3.5 Result

Table 1 presents the first estimation results. For the PC model, the parameter estimates of γ_p and δ_p are of correct signs in accordance with theory (13) and are significant. The estimate of $\gamma_p + \delta_p$ is also of correct sign, but is not significant. According to these estimates, the weight of welfare in the government's objective ($\beta \equiv \frac{a}{1+a}$) is 0.9890, while the fraction of population represented by a lobby (α_L) is 0.7981. The results are broadly consistent with those of Goldberg and Maggi (1999), and serve as the benchmark against which to evaluate the HC model. For the HC model, the clustering of sectors is based on the simplistic criterion, whereby a sector is considered monopolistically competitive if its elasticity measure is greater than one and perfectly competitive otherwise. By this criterion, there are 59 monopolistically competitive sectors and 47 perfectly competitive sectors. The parameter estimates of γ_p and δ_p (and the implied γ_m and δ_m) are of correct signs, and the estimate of $\gamma_p + \delta_p$ is greater than zero (it follows that the estimate of $\gamma_m + \delta_m$ is greater than one); all are in accordance with theories (13) and (14). They are, however, not statistically significant (except for γ_m). This may be due to the fact that the degree of freedom is greatly reduced with the sample broken into two regimes. According to the estimates of the HC model, the weight of welfare in the government's objective ($\beta \equiv \frac{a}{1+a}$) is higher at 0.9988, while the fraction of population represented by a lobby (α_L) is lower at 0.4545. The relatively higher weight of welfare in the government's objective implied by the HC model may be explained by the fact that the welfare-maximizing tariff for a monopolistically competitive sector is positive.[1] Thus, even if a government were not politically motivated (with $\beta = 1$), a monopolistically competitive sector would still receive a positive level of protection. Thus, for a given level of observed protection, it would imply a higher weight of political contribution (and a lower weight of welfare) placed by the government if the sector is classified as perfectly competitive than if it is classified as monopolistically competitive. Treating all sectors as perfectly competitive causes the parameter

[1] The welfare-maximizing tariff for a monopolistically competitive sector is implied by (14) with $a \to \infty$.

estimate of a to bias downward.

Next, the J-test results in Table 1 indicate that neither the PC model nor the HC model with the simplistic clustering criterion is a correct specification of the underlying protection structure. Either model when considered as the null hypothesis is rejected at the 5% significance level. This negative result points to two potential directions of investigation. First, neither theoretical model is complete; important structures or variables are omitted. Second, the market structure dummy under the HC model is not correctly specified and there is room for improvement. We leave the first task to future research and investigate the second possibility below. Indeed, the summary statistics in Table 1 indicate that the derived market structure dummy I^m based on the simplistic criterion is less than satisfactory. The average degree of seller competition is higher in monopolistically competitive sectors (0.2466) than in perfectly competitive sectors (0.2283), when in theory the ranking should be reversed. Thus, there is a need to further refine the market structure dummy.

Toward this end, we adopt the stricter criteria introduced in Section 3.1, that is, $I^m = 1$, if $e_i > 1$ and $scomp_i \leq \bar{\kappa}$. The detailed estimation results for $\bar{\kappa} = \{0.7, 0.6, 0.5, 0.4, 0.3, 0.2\}$ are given in Tables 2 to 7. The threshold levels listed above represent respectively the 94%, 92%, 88%, 82%, 76%, and 62% cutoff of the observed values of '*scomp*'. As can be seen from Tables 2 to 7, the number of sectors classified as monopolistically competitive decreases as the cutoff level of seller competition decreases; the average degree of seller competition for monopolistically competitive sectors naturally decreases as well. With $\bar{\kappa} = 0.7$ in Table 2, monopolistically competitive sectors have already shown a theoretically-desired lower degree of seller competition than perfectly competitive sectors. In addition, the data also seem to suggest a systematic trend for $\bar{\kappa}$ between 0.7 and 0.3: as the criterion becomes more stringent ($\bar{\kappa} \downarrow$), sectors classified as monopolistically competitive have a larger import demand elasticity on average ($\bar{e} \uparrow$), have a higher tendency to become politically organized ($\bar{I} \uparrow$), and receive a higher average protection level ($\bar{t} \uparrow$).

The PC model is independent of the market structure dummy I^m; thus, the estimates of the PC model remain the same across alternative sector classifications in Tables 2 to 7. The estimates of the HC model do not vary significantly for $\bar{\kappa}$ between 0.7 and 0.3: the parameter estimates of γ_p and δ_p (and the implied estimates of γ_m and δ_m) and their sums are of correct signs, but remain statistically insignificant (with the exception of γ_m). As the condition becomes too stringent for a sector to be classified as monopolistically competitive, few monopolistically competitive sectors are left under the HC model, rendering it practically akin to the PC model. This is represented by the case of $\bar{\kappa} = 0.2$ in Table 7, where we see a major shift of the structural estimates of the HC model toward those of the PC model.

The J-test results in Tables 2 to 7 still reject both models, except in Table 6, where the HC model is accepted by setting $\bar{\kappa} = 0.3$ for the degree of seller competition in the construction of the market structure dummy. Also notice that the likelihood of the HC model rises initially as $\bar{\kappa}$ decreases and falls after reaching some maximum. In particular, the likelihood of the accepted HC model with $\bar{\kappa} = 0.3$ is the largest among $\bar{\kappa} = \{0.7, 0.6, 0.5, 0.4, 0.3, 0.2\}$. Thus, we may claim that the HC model coupled with this particular market structure classification provides the best fit so far to the data and is favored by the J-test over the PC model.

To verify the robustness of the above results, we conduct the J-test sequentially for an extensive range of $\bar{\kappa}$, starting with the largest observed value of seller competition (1.3865) and then the next largest, and so forth. We continue lowering the cutoff level and reducing the potential pool of monopolistically competitive sectors this way, until only 35 sectors are left that are classified as monopolistically competitive. This corresponds to $\bar{\kappa} = 0.1873$. The lower bound on the number of classified monopolistically competitive sectors is chosen based on the considerations that there are 22 parameters to be estimated for each market structure in the reduced form equations for z and w, and that some reasonable degree of freedom has to be maintained. Moreover, as discussed above, when the number of sectors classified as monopolistically competitive becomes too small, the HC model approaches the PC model; thus, even if the PC model is accepted against the HC model under such classifications, the finding is not so meaningful. Table 8 presents the summary of the J-test results for the range of $\bar{\kappa}$ that we explored. As can be seen from the table, the likelihood of the HC model tends to rise initially (albeit not strictly monotonically) when the cutoff level decreases, and then to fall back again as the cutoff level becomes too low. The maximum likelihood of the HC model occurs at $\bar{\kappa} = 0.2513$ when the number of monopolistically competitive sectors equals 41. Given the HC model estimates with various $\bar{\kappa}$, the PC model is consistently rejected; on the other hand, the HC model is accepted for a small range of $\bar{\kappa}$ between 0.3233 and 0.2513 with the number of monopolistically competitive sectors ranging between 45 and 41. Based on the HC model with the market structure classification that yields the largest likelihood, the weight of welfare in the government's objective ($\beta \equiv \frac{a}{1+a}$) is estimated at 0.9989 and the fraction of population politically represented (α_L) at 0.5856.

These results suggest a higher weight on welfare placed by the government (β) than indicated by previous studies in the literature based on the G-H homogeneous perfect competition model (0.986 in Goldberg and Maggi, 1999; 0.96 in Eicher and Osang, 2002). This is consistent with our earlier discussion that misspecifying a monopolistically competitive sector as perfectly competitive tends to assign the supposedly welfare-driven component of protection as politically motivated, and to bias upward the estimate of the government's weight on political contribution or to bias downward the government's emphasis on aggregate welfare. The estimates of the degree of political representation (α_L) vary a lot in the literature, ranging from 0.98 in Gawande and Bandyopadhyay (2000), 0.88 in Goldberg and Maggi (1999), to 0.26 in Eicher and Osang (2002). Our estimate of α_L at 0.5856 suggests an intermediate degree of political participation.

4. CONCLUSION

Previous cross-sectional empirical works based on Grossman and Helpman (1994) have often adopted homogeneous (perfect competition) market structure, where all sectors are deemed to be perfectly competitive and share the same protection structure according to G-H. In this paper, we adopt a general empirical specification that accommodates both monopolistically and perfectly competitive sectors, with the protection structure for the monopolistically competitive sectors guided by the recent theory of Chang (2005). The results of the paper cast doubt on the homogeneous perfect competition hypothesis and suggest a direction of improvement toward the heterogeneous market structure hypothesis. Thus, the Chang (2005) model is borne out by the data for monopolistically competitive sectors, while the predictions of G-H remain valid for the subset of perfectly competitive sectors. Our empirical finding thus suggests an intricate relationship between import protection and import penetration: it depends on whether the sector is perfectly competitive or monopolistically competitive and whether the sector is politically organized or not. When all these different facets of heterogeneity are taken into account, the resulting estimates imply that the government's weight on aggregate welfare is higher than previously suggested by the literature (with similar estimation framework) and is close to being a welfare maximizer. On the other hand, the degree of political representation falls in the intermediate range; around half the population is politically represented in the trade policy arena.

A caution is warranted when one would like to take the results given here and stand to applaud the benevolence of the U.S. government in their trade policy making, for one important sector is missing in the sample: the agriculture sector. The sample studied here and in previous G-H-type empirical papers has focused mainly on manufacturing sectors where the general protection level is low, and has excluded the agriculture sector where heavy protection persists. With the agriculture sector included, the conclusions are likely to change toward finding a more special-interest driven government.

APPENDIX A: ESTIMATION METHODOLOGY FOR TOBIT MODEL WITH ENDOGENOUS REGRESSORS

In this appendix, we describe the estimators we explored in our preliminary analysis. The notations here differ somewhat from those in the main text for the sake of simplification. This appendix will also serve as a review on the possible estimation methods employed in previous G-H type empirical studies. Define an indicator function $1[A] = 1$ if A holds and 0 otherwise. Let SF and RF stand for 'structural form' and 'reduced form', respectively. Suppose the equations under consideration are, with $c > 0$,

$$\begin{aligned}
y \text{ SF} &: \quad y_i = \max\{0, \min(y_i^*, c)\}, \quad \text{where } y^* \text{ SF is } \; y_i^* = \beta_w w_i + \beta_{dw} d_i w_i + u_i, \\
w_i &= z_i' \eta_w + v_{wi}, \quad \text{where } v_{wi} \text{ and } u_i \text{ are dependent} \\
d_i &= 1[d_i^* > 0], \quad \text{where } d_i^* = z_i' \eta_d + v_{di}, \;\; v_{di} \text{ and } u_i \text{ are dependent,} \qquad \text{(Model 1)} \\
\text{observed} &: \quad d_i, \; z_i, \; w_i, \; y_i, \quad i = 1, ..., N, \;\; iid \text{ across } i.
\end{aligned}$$

The goal is to estimate the SF parameters β_w and β_{dw}. The upper censoring point c can be allowed to vary across i so long as c_i is observed and independent of y_i. A simpler version of Model 1 is obtained if d is exogenous:

$$\text{same as Model 1 but } d \text{ is exogenous under independence between } v_d \text{ and } u. \qquad \text{(Model 2)}$$

In this case, the slope of w in $y^* = (\beta_w + \beta_{dw} d)w + u$ shifts exogenously depending on d. Substituting the w equation into the y^* SF, we get the y^* RF:

$$y^* \text{ RF} : y^* = z'(\beta_w \eta_w) + dz'(\beta_{dw} \eta_w) \; + \{(\beta_w + \beta_{dw} d)v_w + u\}$$

where the error term in $\{\cdot\}$ is heteroscedastic depending on d and consists of two errors.

Write the y^* RF simply as $y_i^* = x_i'\alpha + \varepsilon_i$, such that $y_i = \max\{0, \min(x_i'\alpha + \varepsilon_i, c)\}$, where x_i is the exogenous regressor vector, α is the parameter vector, and ε_i is the error term. Under the independence of ε from x and the normality $\varepsilon \sim N(0, \sigma_\varepsilon^2)$, the *Censored MLE (CMLE)* is obtained with the log-likelihood function

$$\sum_i \{1[y_i = 0] \ln \Phi(\frac{-x_i'\alpha}{\sigma_\varepsilon}) + 1[0 < y_i < c] \ln\{\frac{\phi((y_i - x_i'\alpha)/\sigma_\varepsilon)}{\sigma_\varepsilon}\} + 1[y_i = c] \ln \Phi(\frac{x_i'\alpha - c}{\sigma_\varepsilon})\}$$

which is maximized with respect to (wrt) α and σ_ε. CMLE converges reasonably well.

Denoting the conditional median of $\varepsilon|x$ as $Med(\varepsilon|x)$, a semiparametric estimator requiring only $Med(\varepsilon|x) = 0$ while allowing an unknown form of heteroskedasticity is the *Censored Least Absolute Deviation estimator (CLAD)* of Powell (1984). For the double censoring case, CLAD minimizes wrt α

$$\sum_i |y_i - \max\{0, \min(x_i'\alpha, c)\}|.$$

The requisite assumption for CLAD is weaker than that for CMLE. Also, no 'nuisance parameter' such as σ_ε appears for CLAD.

Eicher and Osang (2002) apply *Minimum Distance Estimator* (MDE) as explained in Lee (1996a). But MDE requires a somewhat different model from above: instead of the d equation, suppose

$$dw \text{ equation} : d_i w_i = z_i' \eta_{dw} + v_{dwi}.$$

Then, substituting the above w equation and this dw equation into the above y^* SF yields

$$\begin{aligned} y^* \text{ RF-MDE} \quad &: \quad y_i^* = \beta_w(z_i'\eta_w + v_{wi}) + \beta_{dw}(z_i'\eta_{dw} + v_{dwi}) + u_i \equiv z_i'\eta_y + v_y, \quad \text{where} \\ v_y \quad &\equiv \quad \beta_w v_{wi} + \beta_{dw} v_{dwi} + u_i \quad \text{and} \\ \text{MDE Restriction} \quad &: \quad \eta_y = \beta_w \eta_w + \beta_{dw} \eta_{dw}. \end{aligned}$$

This y^* RF in turn leads to

$$y \text{ RF-MDE} : y_i = \max\{0, \min(z_i'\eta_y + v_y, c)\}.$$

The MDE proceeds in two steps. First, estimate η_y for the y RF-MDE with a censored model estimator, and η_w and η_{dw} with Least Squares Estimator (LSE); denote the estimators as h_y, h_w, and h_{dw}, respectively. Second, do the LSE of h_y on h_w and h_{dw} to estimate the two scalars β_w and β_{dw}, which works owing to the MDE Restriction above.

The MDE procedure works well in practice. But the shortcoming is the linear model assumption for dw, which is not tenable in principle because dw is not a continuously distributed random variable due to $P(d = 0) = P(dw = 0) > 0$. Recognizing this problem, one may apply a little different MDE: use a censored model

$$d_i w_i = \max(0,\ z_i'\eta_{dw} + v_{dwi})$$

and estimate this with a CMLE. This should provide a better approximation for dw, but leads to a new problem that the MDE restriction—and consequently y RF-MDE—holds only on $d = 1$. Estimating the y RF-MDE using only the subsample $d = 1$ causes the usual sample selection problem if d is endogenous. If d is exogenous, then the MDE procedure works so long as the estimation of the y RF-MDE is done with the $d = 1$ subsample. That is, the MDE procedure is tenable under Model 2.

Lee (1996b) proposes a *Two-Stage-LSE (2SLS)* type approach for limited dependent variable models with endogenous regressors. The idea is to replace each endogenous regressor with its conditional mean given the instruments. For this, rewrite the y^* SF as

$$\begin{aligned} y^* \text{ SF-2SLS-1} \quad : \quad & y_i^* = \beta_w E(w|z_i) + \beta_{dw} E(dw|z_i) \\ & +u_i + \beta_w\{w_i - E(w|z_i)\} + \beta_{dw}\{d_i w_i - E(dw|z_i)\} \end{aligned}$$

where the last three terms constitute the new (composite) error term. The conditional means $E(w|z)$ and $E(dw|z)$ can be estimated nonparametrically. In practice, the LSE for $w_i = z_i'\eta_w + v_{wi}$ and $d_i w_i = z_i'\eta_{dw} + v_{dwi}$ will do. The latter (practical) approach has the same problem as the above

MDE, because the linear model for the product dw is not tenable in principle.

The 2SLS approach can be similarly applied to Model 2. Rewrite the y^* SF as

$$y^* \text{ SF-2SLS-2} \quad : \quad y_i^* = \beta_w E(w|z_i) + \beta_{dw} d_i E(w|z_i)$$
$$+u_i + \beta_w\{w_i - E(w|z_i)\} + \beta_{dw} d_i \{w_i - E(w|z_i)\}$$

where the last three terms constitute the new (composite) error term. The conditional mean $E(w|z)$ can be estimated using the LSE for $w_i = z_i' \eta_w + v_{wi}$. Hence the problem mentioned above wrt y^* RF-MDE and y^* SF-2SLS-1 does not arise in this case. The y^* SF-2SLS leads straightforwardly to the y SF-2SLS, to which CMLE or CLAD can be applied.

Considering our discussions so far, we can think of at least four methods to estimate β_d and β_{dw}:

	CMLE	CLAD
MDE	y RF estimated, normality	y RF estimated, zero median
2SLS	y SF-2SLS estimated, normality	y SF-2SLS estimated, zero median

Other than the approaches in this table, one may also attempt to apply the system MLE to Model 1 under $(u, v_w, v_d) \sim N(0, \Omega)$, where the covariance matrix Ω has $SD(u)$, $SD(v_w)$, $COR(u, v_w)$, $COR(u, v_d)$, and $COR(v_w, v_d)$; $SD(v_d)$ is not identified. But this MLE has two problems. First, with d endogenous, it seems difficult to obtain the likelihood function. Second, even if the likelihood function is found, estimating the correlation coefficients (and the standard deviations) is notoriously difficult in multivariate MLE's. In practice, often some correlations are assumed to be zero (e.g., $COR(v_w, v_d) = 0$), or an equi-correlation assumption is imposed (e.g., $COR(u, v_w) = COR(u, v_d)$). It is not clear how Goldberg and Maggi (1999) proceeded, as the likelihood function and the estimate for Ω for their MLE were not shown. Alternatively, if Model 2 is adopted, the likelihood function becomes straightforward, with the size of the covariance matrix greatly reduced. Overall, in view of the econometrics issues discussed above concerning the methodological problems implementing Model 1, we choose to adopt Model 2 in the main text.

APPENDIX B: ESTIMATION METHODOLOGY FOR THE STANDARD ERRORS OF THE 2SLS-CMLE

Let α be the first stage parameter of dimension $k_1 \times 1$ and a_N be the LSE. Let β be the likelihood parameter of dimension $k_2 \times 1$ for the second stage, and b_N be the MLE. Denote the second stage score function as $s(z_i, \alpha, \beta)$; omit z_i for simplification. Define

$$\underset{k_2 \times k_1}{\nabla_\alpha s(\alpha,\beta)} \equiv \frac{\partial s(\alpha,\beta)}{\partial \alpha'} \quad \text{and} \quad \underset{k_2 \times k_2}{\nabla_\beta s(\alpha,\beta)} \equiv \frac{\partial s(\alpha,\beta)}{\partial \beta'}.$$

By the definition of b_N, it holds that, using Taylor's expansion,

$$\begin{aligned} 0 &= \frac{1}{\sqrt{N}}\sum_i s(a_N, b_N) \\ &\Longrightarrow \quad 0 \simeq \frac{1}{\sqrt{N}}\sum_i s(a_N,\beta) + \{\frac{1}{N}\sum_i \nabla_\beta s(\alpha,\beta)\}\sqrt{N}(b_N-\beta) \\ &\Longrightarrow \quad \sqrt{N}(b_N-\beta) = -\{\frac{1}{N}\sum_i \nabla_\beta s(\alpha,\beta)\}^{-1}\frac{1}{\sqrt{N}}\sum_i s(a_N,\beta) \\ &\Longrightarrow \quad \sqrt{N}(b_N-\beta) = -\{\frac{1}{N}\sum_i s(\alpha,\beta)s(\alpha,\beta)'\}^{-1}\frac{1}{\sqrt{N}}\sum_i s(a_N,\beta). \end{aligned}$$

To account for the first-stage error $a_N - \alpha$, define

$$H_N \equiv \frac{1}{N}\sum_i \nabla_\beta s(\alpha,\beta) = \frac{1}{N}\sum_i s(\alpha,\beta)s(\alpha,\beta)' \quad \text{and} \quad L_N \equiv \frac{1}{N}\sum_i \nabla_\alpha s(\alpha,\beta).$$

Apply Taylor's expansion to $s(a_N,\beta)$ in the above expression for $\sqrt{N}(b_N-\beta)$ to get

$$\sqrt{N}(b_N-\beta) \simeq -H_N^{-1}\cdot\{\frac{1}{\sqrt{N}}\sum_i s(\alpha,\beta) + L_N\sqrt{N}(a_N-\alpha)\}.$$

With r_i denoting the first-stage LSE residual with Z_i as the regressor, observe

$$\begin{aligned} \sqrt{N}(a_N-\alpha) &= \frac{1}{\sqrt{N}}\sum_i(\frac{1}{N}\sum_i Z_iZ_i')^{-1}Z_ir_i \\ &\equiv \frac{1}{\sqrt{N}}\sum_i \eta_i, \quad \text{where } \eta_i \equiv (\frac{1}{N}\sum_i Z_iZ_i')^{-1}Z_ir_i. \end{aligned}$$

Hence

$$\begin{aligned}
\sqrt{N}(b_N-\beta) &\simeq -H_N^{-1}\{\frac{1}{\sqrt{N}}\sum_i s(\alpha,\beta)+L_N\frac{1}{\sqrt{N}}\sum_i \eta_i\} \\
&= -H_N^{-1}\frac{1}{\sqrt{N}}\sum_i\{s(\alpha,\beta)+L_N\eta_i\} \\
&= -H_N^{-1}\frac{1}{\sqrt{N}}\sum_i q_i, \quad \text{where } q_i\equiv s(\alpha,\beta)+L_N\eta_i.
\end{aligned}$$

Therefore, with $\rightsquigarrow$ denoting convergence in law,

$$\sqrt{N}(b_N-\beta)\rightsquigarrow N(0,H^{-1}E(qq')H^{-1}) \quad \text{where } H\equiv E\{s(\alpha,\beta)s(\alpha,\beta)'\}.$$

Consistent estimators for H and $E(qq')$ are

$$\begin{aligned}
\hat{H}_N &\equiv \frac{1}{N}\sum_i s(a_N,b_N)s(a_N,b_N)', \\
Q_N &\equiv \frac{1}{N}\sum_i\{s(a_N,b_N)+\hat{L}_N\eta_i\}\{s(a_N,b_N)+\hat{L}_N\eta_i\}' \quad \text{where } \hat{L}_N\equiv\frac{1}{N}\sum_i\nabla_\alpha s(a_N,b_N).
\end{aligned}$$

In practice, s can be obtained by numerical derivatives, and $\nabla_\alpha s$ can be obtained using numerical derivatives once more. If the first-stage LSE involves two equations, each with regressor Z_{ji}, residual r_{ji}, η_{ji}, and the parameter α_j, $j=1,2$, then set $\alpha\equiv(\alpha_1',\alpha_2')'$, $a_N\equiv(a_{1N}',a_{2N}')'$, $\eta_i\equiv(\eta_{1i}',\eta_{2i}')'$ and proceed as above.

REFERENCES

Chang, P.-L., 2005. Protection for sale under monopolistic competition, forthcoming in Journal of International Economics.

Davidson, R., MacKinnon, J. G., 1981. Several tests for model specification in the presence of alternative hypothesis. Econometrica 49 (3), 781–93.

Dixit, A. K., Stiglitz, J. E., 1977. Monopolistic competition and optimum product diversity. American Economic Review 67 (3), 297–308.

Eicher, T., Osang, T., 2002. Protection for sale: An empirical investigation: Comment. American Economic Review 92 (5), 1702–1710.

Gawande, K., Bandyopadhyay, U., 2000. Is protection for sale? Evidence on the Grossman-Helpman theory of endogenous protection. Review of Economics and Statistics 82 (1), 139–152.

Goldberg, P. K., Maggi, G., 1999. Protection for sale: An empirical investigation. American Economic Review 89 (5), 1135–1155.

Grossman, G. M., Helpman, E., 1994. Protection for sale. American Economic Review 84 (4), 833–850.

Lee, M.-J., 1996a. Methods of moments and semiparametric econometrics for limited dependent variable models. Springer Verlag, New York.

Lee, M.-J., 1996b. Nonparametric two stage estimation of simultaneous equations with limited endogenous regressor. Econometric Theory 12, 305–330.

Powell, J. L., 1984. Least absolute deviations estimation for the censored regression model. Journal of Econometrics 25, 303–325.

Shiells, C. R., Stern, R. M., Deardorff, A. V., 1986. Estimates of the elasticities of substitution between imports and home goods for the United States. Weltwirtschaftliches Archiv 122 (3), 497–519.

Trefler, D., 1993. Trade liberalization and the theory of endogenous protection. Journal of Political Economy 101 (1), 138–160.

Table 1: clustering of sectors by elasticity ($e_i > 1$)

	PC		HC	
γ_p	-0.0088***	(0.0032)	-0.0005	(0.0005)
δ_p	0.0110*	(0.0059)	0.0012	(0.0010)
σ_{ϵ_p}	0.6693	(0.1037)	0.1219	(0.0144)
$\rho_{u_p,\epsilon_p}(\rho_{u_{pp},\epsilon_p})$	-0.9149	(0.6486)	-1	
γ_m			0.9995***	(0.0005)
δ_m			0.0012	(0.0010)
σ_{ϵ_m}			1.3093	(0.0432)
ρ_{u_{mm},ϵ_m}			-1	
$\gamma_p + \delta_p$	0.0022	(0.0045)	0.0007	(0.0010)
$\gamma_m + \delta_m - 1$			0.0007	(0.0010)
a	90.1326*	(48.1782)	849.0493	(740.7607)
α_L	0.7981**	(0.3271)	0.4545	(0.4453)
$L_p(L_h)$	-99.0480		-78.3968	
J test:				
$\mu_h(\mu_p)$	0.6914**	(0.2860)	0.4271**	(0.1841)
$\tilde{L}_p(\tilde{L}_h)$	-94.2304		-76.0305	

Summary Statistics:	# of sectors	$\bar{e}$	$sc\bar{o}mp$	$\bar{I}$	$\bar{t}$
$I_i^m = 1$	59	4.0337	0.2466	0.3898	0.1203
$I_i^m = 0$	47	0.4492	0.2283	0.4681	0.1535
Total	106	2.4443	0.2385	0.4245	0.1350

Note:
1. $I_i^m = 1$ if $e_i > 1$. The variable 'seller competition (*scomp*)' is the seller number of firms in an industry scaled by industry sales.
2. Figures in brackets are standard errors.
3. A * sign indicates significance at 10% level, a ** sign indicates significance at 5% level, and a *** sign indicates significance at 1% level.
4. For parameters, σ_{ϵ_p} and σ_{ϵ_m}, which are not tested against zero, we do not attach the * sign.
5. The estimates of ρ_{u_{pp},ϵ_p} and ρ_{u_{mm},ϵ_m} in the HC model reach the lower bound -1. Note that $\rho = -1$ does not make the likelihood function of the HC model degenerate as it would do in a bivariate normal density function.

Table 2: clustering of sectors by elasticity ($e_i > 1$) and seller competition ($scomp_i \leq 0.7$)

	PC		HC	
γ_p	-0.0088***	(0.0032)	-0.0006	(0.0005)
δ_p	0.0110*	(0.0059)	0.0012	(0.0009)
σ_{ϵ_p}	0.6693	(0.1037)	0.1187	(0.0145)
$\rho_{u_p,\epsilon_p}(\rho_{u_{pp},\epsilon_p})$	-0.9149	(0.6486)	-1	
γ_m			0.9994***	(0.0005)
δ_m			0.0012	(0.0009)
σ_{ϵ_m}			1.3038	(0.0433)
ρ_{u_{mm},ϵ_m}			-1	
$\gamma_p + \delta_p$	0.0022	(0.0045)	0.0006	(0.0008)
$\gamma_m + \delta_m - 1$			0.0006	(0.0008)
a	90.1326*	(48.1782)	857.9891	(657.7732)
α_L	0.7981**	(0.3271)	0.4926	(0.4155)
$L_p(L_h)$	-99.0480		-75.1774	
J test:				
$\mu_h(\mu_p)$	0.7024**	(0.2892)	0.4197**	(0.1668)
$\tilde{L}_p(\tilde{L}_h)$	-94.1833		-72.2502	

Summary Statistics:	# of sectors	$\bar{e}$	$sc\bar{o}mp$	$\bar{I}$	$\bar{t}$
$I_i^m = 1$	56	4.0699	0.1991	0.4107	0.1267
$I_i^m = 0$	50	0.6236	0.2825	0.4400	0.1444
Total	106	2.4443	0.2385	0.4245	0.1350

Note:

1. $I_i^m = 1$ if $e_i > 1$ and $scomp_i \leq 0.7$, where the variable 'seller competition ($scomp$)' is the seller number of firms in an industry scaled by industry sales.
2. Figures in brackets are standard errors.
3. A * sign indicates significance at 10% level, a ** sign indicates significance at 5% level, and a *** sign indicates significance at 1% level.
4. For parameters, σ_{ϵ_p} and σ_{ϵ_m}, which are not tested against zero, we do not attach the * sign.
5. The estimates of ρ_{u_{pp},ϵ_p} and ρ_{u_{mm},ϵ_m} in the HC model reach the lower bound -1. Note that $\rho = -1$ does not make the likelihood function of the HC model degenerate as it would do in a bivariate normal density function.

Table 3: clustering of sectors by elasticity ($e_i > 1$) and seller competition ($scomp_i \leq 0.6$)

	PC		HC	
γ_p	-0.0088***	(0.0032)	-0.0005	(0.0005)
δ_p	0.0110*	(0.0059)	0.0011	(0.0009)
σ_{ϵ_p}	0.6693	(0.1037)	0.1229	(0.0150)
$\rho_{u_p,\epsilon_p}(\rho_{u_{pp},\epsilon_p})$	-0.9149	(0.6486)	-1	
γ_m			0.9995***	(0.0005)
δ_m			0.0011	(0.0009)
σ_{ϵ_m}			1.3335	(0.0454)
ρ_{u_{mm},ϵ_m}			-1	
$\gamma_p + \delta_p$	0.0022	(0.0045)	0.0006	(0.0008)
$\gamma_m + \delta_m - 1$			0.0006	(0.0008)
a	90.1326*	(48.1782)	909.7454	(760.2341)
α_L	0.7981**	(0.3271)	0.4318	(0.4343)
$L_p(L_h)$	-99.0480		-72.7223	
J test:				
$\mu_h(\mu_p)$	0.7049**	(0.2937)	0.3969**	(0.1622)
$\tilde{L}_p(\tilde{L}_h)$	-94.1821		-69.8794	

Summary Statistics:	# of sectors	$\bar{e}$	$sc\bar{o}mp$	$\bar{I}$	$\bar{t}$
$I_i^m = 1$	53	4.1867	0.1741	0.4340	0.1280
$I_i^m = 0$	53	0.7019	0.3028	0.4151	0.1421
Total	106	2.4443	0.2385	0.4245	0.1350

Note:

1. $I_i^m = 1$ if $e_i > 1$ and $scomp_i \leq 0.6$, where the variable 'seller competition ($scomp$)' is the seller number of firms in an industry scaled by industry sales.
2. Figures in brackets are standard errors.
3. A * sign indicates significance at 10% level, a ** sign indicates significance at 5% level, and a *** sign indicates significance at 1% level.
4. For parameters, σ_{ϵ_p} and σ_{ϵ_m}, which are not tested against zero, we do not attach the * sign.
5. The estimates of ρ_{u_{pp},ϵ_p} and ρ_{u_{mm},ϵ_m} in the HC model reach the lower bound -1. Note that $\rho = -1$ does not make the likelihood function of the HC model degenerate as it would do in a bivariate normal density function.

Table 4: clustering of sectors by elasticity ($e_i > 1$) and seller competition ($scomp_i \leq 0.5$)

	PC		HC	
γ_p	-0.0088***	(0.0032)	-0.0007	(0.0005)
δ_p	0.0110*	(0.0059)	0.0014	(0.0010)
σ_{ϵ_p}	0.6693	(0.1037)	0.1326	(0.0170)
$\rho_{u_p,\epsilon_p}(\rho_{u_{pp},\epsilon_p})$	-0.9149	(0.6486)	-1	
γ_m			0.9993***	(0.0005)
δ_m			0.0014	(0.0010)
σ_{ϵ_m}			1.3507	(0.0489)
ρ_{u_{mm},ϵ_m}			-1	
$\gamma_p + \delta_p$	0.0022	(0.0045)	0.0007	(0.0009)
$\gamma_m + \delta_m - 1$			0.0007	(0.0009)
a	90.1326*	(48.1782)	696.3284	(472.3664)
α_L	0.7981**	(0.3271)	0.4718	(0.3636)
$L_p(L_h)$	-99.0480		-71.3375	
J test:				
$\mu_h(\mu_p)$	0.7226**	(0.2975)	0.4307***	(0.1630)
$\tilde{L}_p(\tilde{L}_h)$	-93.8804		-68.1527	

Summary Statistics:	# of sectors	$\bar{e}$	$sc\bar{o}mp$	$\bar{I}$	$\bar{t}$
$I_i^m = 1$	50	4.3010	0.1500	0.4400	0.1320
$I_i^m = 0$	56	0.7866	0.3175	0.4107	0.1378
Total	106	2.4443	0.2385	0.4245	0.1350

Note:

1. $I_i^m = 1$ if $e_i > 1$ and $scomp_i \leq 0.5$, where the variable 'seller competition ($scomp$)' is the seller number of firms in an industry scaled by industry sales.
2. Figures in brackets are standard errors.
3. A * sign indicates significance at 10% level, a ** sign indicates significance at 5% level, and a *** sign indicates significance at 1% level.
4. For parameters, σ_{ϵ_p} and σ_{ϵ_m}, which are not tested against zero, we do not attach the * sign.
5. The estimates of ρ_{u_{pp},ϵ_p} and ρ_{u_{mm},ϵ_m} in the HC model reach the lower bound -1. Note that $\rho = -1$ does not make the likelihood function of the HC model degenerate as it would do in a bivariate normal density function.

Table 5: clustering of sectors by elasticity ($e_i > 1$) and seller competition ($scomp_i \leq 0.4$)

	PC		HC	
γ_p	-0.0088***	(0.0032)	-0.0006	(0.0005)
δ_p	0.0110*	(0.0059)	0.0012	(0.0010)
σ_{ϵ_p}	0.6693	(0.1037)	0.1277	(0.0156)
$\rho_{u_p,\epsilon_p}(\rho_{u_{pp},\epsilon_p})$	-0.9149	(0.6486)	-1	
γ_m			0.9994***	(0.0005)
δ_m			0.0012	(0.0010)
σ_{ϵ_m}			1.3710	(0.0556)
ρ_{u_{mm},ϵ_m}			-1	
$\gamma_p + \delta_p$	0.0022	(0.0045)	0.0006	(0.0009)
$\gamma_m + \delta_m - 1$			0.0006	(0.0009)
a	90.1326*	(48.1782)	807.7733	(629.8685)
α_L	0.7981**	(0.3271)	0.4926	(0.4354)
$L_p(L_h)$	-99.0480		-65.1257	
J test:				
$\mu_h(\mu_p)$	0.7376**	(0.2966)	0.3705**	(0.1478)
$\tilde{L}_p(\tilde{L}_h)$	-93.5365		-62.2300	

Summary Statistics:	# of sectors	$\bar{e}$	$sc\bar{o}mp$	$\bar{I}$	$\bar{t}$
$I_i^m = 1$	47	4.3151	0.1314	0.4681	0.1383
$I_i^m = 0$	59	0.9540	0.3238	0.3898	0.1325
Total	106	2.4443	0.2385	0.4245	0.1350

Note:

1. $I_i^m = 1$ if $e_i > 1$ and $scomp_i \leq 0.4$, where the variable 'seller competition ($scomp$)' is the seller number of firms in an industry scaled by industry sales.
2. Figures in brackets are standard errors.
3. A * sign indicates significance at 10% level, a ** sign indicates significance at 5% level, and a *** sign indicates significance at 1% level.
4. For parameters, σ_{ϵ_p} and σ_{ϵ_m}, which are not tested against zero, we do not attach the * sign.
5. The estimates of ρ_{u_{pp},ϵ_p} and ρ_{u_{mm},ϵ_m} in the HC model reach the lower bound -1. Note that $\rho = -1$ does not make the likelihood function of the HC model degenerate as it would do in a bivariate normal density function.

Table 6: clustering of sectors by elasticity ($e_i > 1$) and seller competition ($scomp_i \leq 0.3$)

	PC		HC	
γ_p	-0.0088***	(0.0032)	-0.0006	(0.0004)
δ_p	0.0110*	(0.0059)	0.0011	(0.0008)
σ_{ϵ_p}	0.6693	(0.1037)	0.1206	(0.0152)
$\rho_{u_p,\epsilon_p}(\rho_{u_{pp},\epsilon_p})$	-0.9149	(0.6486)	-1	
γ_m			0.9994***	(0.0004)
δ_m			0.0011	(0.0008)
σ_{ϵ_m}			1.4002	(0.0609)
ρ_{u_{mm},ϵ_m}			-1	
$\gamma_p + \delta_p$	0.0022	(0.0045)	0.0005	(0.0007)
$\gamma_m + \delta_m - 1$			0.0005	(0.0007)
a	90.1326*	(48.1782)	903.2278	(637.4129)
α_L	0.7981**	(0.3271)	0.5463	(0.4303)
$L_p(L_h)$	-99.0480		-57.5475	
J test:				
$\mu_h(\mu_p)$	0.7157**	(0.2792)	0.5024	(0.5322)
$\tilde{L}_p(\tilde{L}_h)$	-93.4472		-54.0826	

Summary Statistics:	# of sectors	$\bar{e}$	$\bar{scomp}$	$\bar{I}$	$\bar{t}$
$I_i^m = 1$	43	4.5025	0.1119	0.4884	0.1489
$I_i^m = 0$	63	1.0395	0.3249	0.3810	0.1256
Total	106	2.4443	0.2385	0.4245	0.1350

Note:

1. $I_i^m = 1$ if $e_i > 1$ and $scomp_i \leq 0.3$, where the variable 'seller competition ($scomp$)' is the seller number of firms in an industry scaled by industry sales.
2. Figures in brackets are standard errors.
3. A * sign indicates significance at 10% level, a ** sign indicates significance at 5% level, and a *** sign indicates significance at 1% level.
4. For parameters, σ_{ϵ_p} and σ_{ϵ_m}, which are not tested against zero, we do not attach the * sign.
5. The estimates of ρ_{u_{pp},ϵ_p} and ρ_{u_{mm},ϵ_m} in the HC model reach the lower bound -1. Note that $\rho = -1$ does not make the likelihood function of the HC model degenerate as it would do in a bivariate normal density function.

Table 7: clustering of sectors by elasticity ($e_i > 1$) and seller competition ($scomp_i \leq 0.2$)

	PC		HC	
γ_p	-0.0088***	(0.0032)	-0.0015**	(0.0006)
δ_p	0.0110*	(0.0059)	0.0023**	(0.0009)
σ_{ϵ_p}	0.6693	(0.1037)	0.1637	(0.0177)
$\rho_{u_p,\epsilon_p}(\rho_{u_{pp},\epsilon_p})$	-0.9149	(0.6486)	-1	
γ_m			0.9985***	(0.0006)
δ_m			0.0023**	(0.0009)
σ_{ϵ_m}			1.3124	(0.0661)
ρ_{u_{mm},ϵ_m}			-1	
$\gamma_p + \delta_p$	0.0022	(0.0045)	0.0008	(0.0008)
$\gamma_m + \delta_m - 1$			0.0008	(0.0008)
a	90.1326*	(48.1782)	427.5867**	(169.9796)
α_L	0.7981**	(0.3271)	0.6374***	(0.2434)
$L_p(L_h)$	-99.0480		-62.8964	
J test:				
$\mu_h(\mu_p)$	0.7367***	(0.2599)	0.6652***	(0.2109)
$\tilde{L}_p(\tilde{L}_h)$	-92.9147		-58.7263	

Summary Statistics:	# of sectors	$\bar{e}$	$sc\bar{o}mp$	$\bar{I}$	$\bar{t}$
$I_i^m = 1$	36	4.2167	0.0872	0.5556	0.1629
$I_i^m = 0$	70	1.5328	0.3163	0.3571	0.1207
Total	106	2.4443	0.2385	0.4245	0.1350

Note:
1. $I_i^m = 1$ if $e_i > 1$ and $scomp_i \leq 0.2$, where the variable 'seller competition ($scomp$)' is the seller number of firms in an industry scaled by industry sales.
2. Figures in brackets are standard errors.
3. A * sign indicates significance at 10% level, a ** sign indicates significance at 5% level, and a *** sign indicates significance at 1% level.
4. For parameters, σ_{ϵ_p} and σ_{ϵ_m}, which are not tested against zero, we do not attach the * sign.
5. The estimates of ρ_{u_{pp},ϵ_p} and ρ_{u_{mm},ϵ_m} in the HC model reach the lower bound -1. Note that $\rho = -1$ does not make the likelihood function of the HC model degenerate as it would do in a bivariate normal density function.

Table 8: summary of the J-test for an extensive range of seller competition thresholds

sector clustering			HC model estimates					model selection	
cutoff value of *scomp* ($\bar{\kappa}$)	cutoff rank of *scomp*	no. of MC sectors	a		α_L		L_h	PC	HC
1.3865	106	59	849.0493	(740.7607)	0.4545	(0.4453)	-78.3968	R**	R**
1.3222	105	59	849.0493	(740.7607)	0.4545	(0.4453)	-78.3968	R**	R**
1.2246	104	59	849.0493	(740.7607)	0.4545	(0.4453)	-78.3968	R**	R**
1.2014	103	58	830.3585	(669.8679)	0.5210	(0.4391)	-77.9522	R**	R**
0.9717	102	57	810.7296	(616.9224)	0.5327	(0.4179)	-77.3769	R**	R**
0.9716	101	56	857.9891	(657.7732)	0.4926	(0.4155)	-75.1774	R**	R**
0.6550	100	56	857.9891	(657.7732)	0.4926	(0.4155)	-75.1774	R**	R**
0.6397	99	55	858.8432	(659.0478)	0.4914	(0.4155)	-74.9309	R**	R**
0.6289	98	54	851.6470	(673.4271)	0.4464	(0.4140)	-74.7138	R**	R**
0.5953	97	53	909.7454	(760.2341)	0.4318	(0.4343)	-72.7223	R**	R**
0.5896	96	52	874.5466	(695.5136)	0.4379	(0.4181)	-72.4156	R**	R**
0.5628	95	51	661.9317	(433.5991)	0.4787	(0.3519)	-73.5389	R**	R***
0.5429	94	51	661.9317	(433.5991)	0.4787	(0.3519)	-73.5389	R**	R***
0.4926	93	50	696.3284	(472.3664)	0.4718	(0.3636)	-71.3375	R**	R***
0.4868	92	50	696.3284	(472.3664)	0.4718	(0.3636)	-71.3375	R**	R***
0.4619	91	50	696.3284	(472.3664)	0.4718	(0.3636)	-71.3375	R**	R***
0.4380	90	49	727.0918	(510.5308)	0.4718	(0.3789)	-69.7894	R**	R**
0.4262	89	48	763.1615	(555.1879)	0.4624	(0.3892)	-67.4248	R**	R**
0.4220	88	47	807.7733	(629.8685)	0.4926	(0.4354)	-65.1257	R**	R**
0.3668	87	47	807.7733	(629.8685)	0.4926	(0.4354)	-65.1257	R**	R**
0.3549	86	46	802.1698	(630.5880)	0.4672	(0.4306)	-63.3556	R**	R***
0.3507	85	46	802.1698	(630.5880)	0.4672	(0.4306)	-63.3556	R**	R***
0.3233	84	45	825.5517	(670.0232)	0.5107	(0.4711)	-61.0028	R***	A
0.3229	83	44	869.5347	(594.5170)	0.5537	(0.4182)	-59.7069	R***	A
0.3115	82	43	903.2278	(637.4129)	0.5463	(0.4303)	-57.5475	R**	A
0.2973	81	43	903.2278	(637.4129)	0.5463	(0.4303)	-57.5475	R**	A
0.2920	80	43	903.2278	(637.4129)	0.5463	(0.4303)	-57.5475	R**	A
0.2623	79	42	920.0701	(664.4405)	0.5341	(0.4374)	-57.1308	R***	A
0.2594	78	42	920.0701	(664.4405)	0.5341	(0.4374)	-57.1308	R***	A
0.2528	77	42	920.0701	(664.4405)	0.5341	(0.4374)	-57.1308	R***	A
0.2513	76	41	888.3396	(593.6911)	0.5856	(0.4136)	-55.9352	R***	A
0.2382	75	40	494.6710	(237.9385)	0.5536	(0.2767)	-67.8295	R***	R***
0.2353	74	39	443.6104	(184.3986)	0.6248	(0.2547)	-66.7684	R**	R***
0.2335	73	39	443.6104	(184.3986)	0.6248	(0.2547)	-66.7684	R**	R***
0.2332	72	38	435.6234	(172.8254)	0.6088	(0.2406)	-64.9324	R***	R***
0.2125	71	38	435.6234	(172.8254)	0.6088	(0.2406)	-64.9324	R***	R***
0.2105	70	38	435.6234	(172.8254)	0.6088	(0.2406)	-64.9324	R***	R***
0.2029	69	38	435.6234	(172.8254)	0.6088	(0.2406)	-64.9324	R***	R***
0.2011	68	37	432.2619	(172.2964)	0.6056	(0.2401)	-63.8470	R***	R***
0.2010	67	36	427.5867	(169.9796)	0.6374	(0.2434)	-62.8964	R***	R***
0.1885	66	36	427.5867	(169.9796)	0.6374	(0.2434)	-62.8964	R***	R***
0.1873	65	35	416.3748	(165.2915)	0.6419	(0.2449)	-62.4645	R***	R***

Note:
1. The variable 'seller competition (*scomp*)' is the seller number of firms in an industry scaled by industry sales. A sector is classified as monopolistically competitive (MC) if $e_i > 1$ and $scomp_i \leq \bar{\kappa}$.
2. Figures in brackets are standard errors.
3. R* indicates rejection at the 10% significance level, R** indicates rejection at the 5% significance level, and R*** indicates rejection at the 1% significance level. 'A' indicates acceptance at least at the 10% significance level.