QUALITY CHANGE UNDER TRADE RESTRAINTS IN
JAPANESE AUTOS*

ROBERT C. FEENSTRA

In this paper we investigate the quality change in Japanese car and truck imports over 1979–1985. Car imports have been subject to a quota restraint since April 1981, while compact trucks have faced an ad valorem tariff of 25 percent since August 1980. We find evidence of substantial upgrading in Japanese car imports, with ambiguous quality change in trucks. The welfare cost of the quota restraint in cars exceeds $1,000 per import in 1983 and 1984.

I. INTRODUCTION

Recent years have seen an increased use of quota restrictions in trade between industrialized countries, rather than ad valorem tariffs. An expected response to a quota on the number of units sold is that a firm may upgrade its product, through changing the design, adding extra features, etc. In this paper we shall test the hypothesis of quality upgrading for U. S. imports of Japanese cars, which were subject to a strict quota beginning in April 1981. Quality change is measured using hedonic regressions [Griliches, 1971]. Our results are contrasted with U. S. imports of Japanese compact trucks, which have been subject to a tariff of 25 percent since August 1980.

The upgrading of imports is of interest because it demonstrates a margin of substitution that can be exploited by firms, bypassing the government regulations. Baldwin [1982] has emphasized this phenomenon in his study of the “inefficacy of trade policy.” From a consumer viewpoint the costs of the trade restraint are substantially affected by upgrading. The nominal rise in the import price under a quota overstates the consumer cost, since part of the price rise is caused by upgrading. A secondary purpose of this paper is to isolate the “pure” price increase in Japanese car imports due to the quota, after correcting for upgrading.

The early theoretical work on upgrading considered consumers buying several exogenous varieties of a good, in which case a quota

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or specific tax very likely caused upgrading. Subsequent research has recognized that quality can be varied continuously and is a strategic choice variable of firms. These models can lead to ambiguous affects of a quota or specific tax. However, existing models often consider just one or several varieties produced in equilibrium. Since we are considering a market with many available varieties, and since our empirical work uses hedonic regressions, we shall adopt the hedonic price model of Rosen [1974] with a continuum of product types. In Section II we examine the impact of quotas and tariffs on the choice of quality by firms. We find that a reduction in quantity sold due to the quota can lead firms to substitute toward higher quality, but this result does not hold for all cost functions. The same incentive for quality upgrading is not present with an ad valorem tariff.

In Section III we briefly discuss the trade policy in Japanese cars and trucks and outline our data. Estimation is performed in Section IV. The hedonic regression for cars must include coefficients allowing prices to rise proportionally over time (due to changes in costs or exchange rates), and must also permit prices to rise by specific dollar amounts over time (this is the pure price effect of the quota). However, multicollinearity prevents an accurate estimate of both sets of coefficients. We overcome this problem by pooling our car and truck data, which permits accurate estimates of the specific price effect of the quota.

Our results are presented in Section V. We find substantial upgrading in Japanese car imports, with ambiguous quality change in trucks. One half of the nominal increase in car prices over 1980–1985 is explained by quality improvement. The pure price effect of the quota exceeds $1,000 per import in 1983 and 1984, with a standard error of $270. It appears that the pricing pattern of Japanese cars has significantly changed in 1985, which may be due to more collusive behavior.

II. HEDONIC PRICE MODEL

The model of Rosen [1974] has consumers and firms choosing their optimal positions along an equilibrium price schedule $p(z)$,

2. See Borcherding and Silberberg [1978] and Falvey [1979].
where \( z \) is a vector of characteristics of the product in question. Consumers purchase one unit of the differentiated product and obtain utility \( U(z, x; \alpha) \), where \( x \) is the quantity of a numeraire good, and \( \alpha \) is a vector of consumer parameters reflecting tastes. Maximizing utility subject to the budget constraint \( y = p(z) + x \), where \( y \) is income, we obtain the first-order conditions:

\[
(1) \quad p_z(z) = U_z[z, y - p(z); \alpha]/U_z[z, y - p(z); \alpha].
\]

The costs of domestic or foreign firms are \( C(M, z; \beta) \), where \( M \) is the quantity produced of the differentiated product with characteristics \( z \), and \( \beta \) is a vector of firm parameters. These parameters reflect firm-specific technological knowledge, as well as differences in factor prices across countries. Firms are competitive and take prices \( p(z) \) as given. Foreign firms face a quota constraint \( M \leq \bar{M} \), where \( \bar{M} \) may differ across firms. The Lagrangian for foreign firms is \( L = p(z)M - C(M, z; \beta) + s(\bar{M} - M) \), where \( s \geq 0 \) is the shadow price of the quota constraint. When the constraint is binding, the first-order conditions for foreign firms are

\[
(2) \quad p_z(z) = C_z(\bar{M}, z; \beta)/\bar{M},
\]

\[
(3) \quad p(z) = C_M(\bar{M}, z; \beta) + s.
\]

Given \( \bar{M} \), (2) determines the optimal choice of \( z \) for a foreign firm. Then (3) determines the value of \( s \), which is the quota rent per unit produced. The full equilibrium conditions are (1), (2), (3), and analogous conditions for domestic firms (with \( s = 0 \) and \( M \) endogenous); and supply equals demand for each product type. Examples of how to determine the equilibrium price schedule \( p(z) \), in the absence of quotas, are provided by Rosen [1974] and Epplle [1987].

A. Effect of the Quota

Suppose that the import quota is tightened, so that \( \bar{M} \) is reduced across foreign firms. This will have two effects on the optimal choice of characteristics. First, the reduction \( \bar{M} \) will change \( C_z \) and directly affect \( z \) from (2). Second, the tightening of the quota will change the equilibrium price schedule \( p(z) \), which can also affect the choice of \( z \) in (2). We shall initially examine the first of these effects.

Differentiating (2) treating the price function \( p(z) \) as fixed, we obtain

\[
(4) \quad \frac{dz}{d\bar{M}} = \left[ p_{zz} - \left( \frac{C_{zz}}{\bar{M}} \right) \right]^{-1} \left[ \frac{C_z M}{\bar{M}} - \left( \frac{C_z}{\bar{M}^2} \right) \right].
\]
The matrix \([p_{zz} - (C_{zz}/\bar{M})]\) and its inverse is negative definite from the second-order conditions for profit maximization. The column vector \([(C_{z\bar{M}}/\bar{M}) - (C_{z\bar{z}}/\bar{M}^2)]\) is the change in the marginal cost of each characteristic when output \(\bar{M}\) varies. Convexity of the cost function in \((M, z)\) does not determine the sign of this vector. To evaluate the overall change in product "quality" due to tightening of the quota, we can premultiply \((4)\) by the row vector \(p'_{z} = C_{z}/\bar{M}\). The resulting expression is of ambiguous sign.

However, intuition suggests that for a reasonable class of cost functions, the marginal cost of some characteristics (or of overall quality) should be decreasing in output. This is because a firm that experiences a decline in output would find itself with unused amounts of fixed inputs, which could be used to upgrade the units being produced. It turns out that this intuition applies for cost functions of the form \(C(M, z; \beta) = c[g(Mz; \beta)] = c[Mg(z; \beta)]\), where \(g\) is homogeneous of degree one and can be thought of as a unit-cost function, and \(c\) is an increasing and convex transformation. This functional form specifies that the relevant units for measuring output are \(Mz\), i.e., the total amount produced of each characteristic.

Using this special form of costs, we can calculate

\[
\frac{p'dz}{dM} = g'_z \left[ p_{zz} - \frac{C_{zz}}{\bar{M}} \right]^{-1} g_z g' c'' \leq 0,
\]

where a strict inequality holds when \(c'' > 0\). Thus, a reduction in output due to the quota will raise product quality. While we have obtained this as the optimal response of foreign firms, a similar result applies to consumers when they purchase more than one unit. Then a specific rise in the import price schedule can cause consumers to reduce quantity purchased and raise quality, but this result does not hold for all utility functions.\(^5\)

We next consider the effect of the quota on the equilibrium price schedule. In the short run the schedule \(p(z)\) could change nonlinearly as firms move along their marginal cost curves, adjusting to the new consumer demands (see Bond [1986]). To simplify the analysis, we focus on the long-run equilibrium [Rosen, 1974, sec. IIIB]. Plants are constructed to achieve minimum average cost,

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\(^5\) This result can be obtained in Rosen's model when utility is of the form \(U(m, z, x; \omega) = u[f(Mz; \omega)] + x = u[Mf(z; \omega)] + x\), where \(m\) is the number of units purchased, \(f\) is homogeneous of degree one, and \(u' > 0, u'' < 0\). The sensitivity of upgrading to the utility function is especially apparent in Leffler [1982] and Krishna [1987].
which is \( h(z; \beta) = \min_M C(M, z; \beta)/M \). Total costs are \( Mh(z; \beta) \) and the first-order conditions of foreign firms become

\[
(2') \quad p_z(z) = h_z(z; \beta),
\]

\[
(3') \quad p(z) = h(z; \beta) + s.
\]

Foreign entry takes the form of firms switching product types within their output quotas, where we assume that in the long run firms have access to the same technology. This implies an equalization of the quota rents earned per unit across foreign firms. The equilibrium foreign price schedule is

\[
(6) \quad p(z) = \phi(z) + s,
\]

where \( \phi(z) = \min_{\beta} h(z; \beta) \) is the envelope of firm’s minimum average cost.

A tightening of the quota in the long run implies a rise in the quota rent \( s \), equal across foreign firms. This acts as a specific price increase in (6). The choice of product characteristics by firms is not affected by tightening the quota, since output \( M \) does not appear in \((2')\) and the derivatives \( p_z = \phi_z \) are unchanged in \((6)\) and \((2')\). In the long run, firms adjust to the quota simply by reducing output and capacity, with constant marginal and average costs.

While there is no long-run effect on quality from firms, consumers can still respond in several ways: by reducing quantity purchased and raising quality (see the discussion following \((5)\)); by switching from imported to domestic varieties; or by exiting the market. The first of these effects would raise the average quality of imports, obtained by multiplying \( p_z \) by the share of demand and integrating over all imports produced. However, the second and third effects have ambiguous effects on average quality, since consumers switching and exiting may have formerly purchased either high or low quality imports (see Feenstra [1986]).

In our study of Japanese auto imports, we shall consider the response of both firms and consumers to the quota restraint. After briefly examining the effect of an ad valorem tariff below, we turn to a discussion of our data.

**B. Effect of an Ad Valorem Tariff**

Suppose that foreign firms are faced with an ad valorem tariff of \( \tau \) rather than the import quota. Profits become \( \pi = p(z)M/\)

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6. Krishna [1985] shows that the optimal response to a quota by a monopolist supplying a continuum of goods is to raise each price by a specific amount, as in (6). Time-series evidence on whether U. S. auto prices correspond to a competitive or monopolistic model is presented in Blanchard and Melino [1984].
(1 + τ) \cdot C(M, z; \beta), where \( p(z) \) is the consumer price schedule. The first-order conditions are

(7) \quad \frac{p_z(z)}{(1 + \tau)} = C_z(M, z; \beta) / M,

(8) \quad \frac{p(z)}{(1 + \tau)} = C_M(M, z; \beta).

As with the quota, we shall examine the effect of tightening the tariff, holding prices \( p(z) \) constant. Performing the comparative statics on (7) and (8), we find that \( dz/d\tau \) is ambiguous. Adopting the special cost function \( C(M, z; \beta) = c[g(Mz; \beta)] = c[Mg(z, \beta)] \), we find that \( dz/d\tau \) is zero. To see this, write the first-order conditions for this cost function as \( p_z/(1 + \tau) = c'g_z \) and \( p/(1 + \tau) = c'g \), which imply that

(9) \quad \frac{p_z(z)}{p(z)} = g_z(z; \beta) / g(z; \beta).

The optimal choice of characteristics is determined by (9), independent of output or the tariff. We can interpret (9) as the condition for minimizing the “average cost of quality,” that is min, \( g(z; \beta)/p(z) \). Changing prices by a scalar multiple does not affect this decision. For the special cost function above, a higher ad valorem tariff leads to reduced output but no change in product quality by firms.\(^8\)

III. JAPANESE CAR AND TRUCK IMPORTS

The legislative history of the quota on Japanese cars is summarized in Feenstra [1984, 1985]. For the April–March periods of 1981, 1982, and 1983, Japanese sales were limited to a total of 1,832,500 units as follows: 1.68 million passenger cars to the United States; 82,500 “utility” vehicles to the United States such as the Subaru Brat, Toyota Land Cruisers, and Van; and 70,000 units exported to Puerto Rico. The quotas are allocated to Japanese firms by their government. For April 1984–March 1985 the restraints were continued and raised to 2.016 million units. Following this, the Japanese government has continued the restraint each year at 2.506 million annually for April 1985–March 1988, of which 2.3 million are passenger car sales to the United States. These extensions were

\(^7\) This result is also established by Rodríguez [1979]. Here \( g(z; \beta) \) is the total cost of producing characteristics \( z \) (and one unit of output), while \( p(z) \) is interpreted as "quality."

\(^8\) The response of consumers to an ad valorem tariff is analyzed in Feenstra [1986].
made amid protectionist sentiments in Congress, but were not formally requested by the United States.

During the 1970s Japan exported an increasing number of compact trucks to the United States, most as cab/chassis with some final assembly needed. These were classified as "parts of trucks" which carried a tariff rate of 4 percent, whereas "complete or unfinished trucks" had a duty of 25 percent. With prodding from Congress, in 1980 the U. S. Customs Service announced that effective August 21 imported lightweight cab/chassis would be reclassified as complete trucks. This raised the tariff rates on nearly all Japanese trucks from 4 to 25 percent.

To determine the impact of the trade restraints on product quality, a sample of Japanese cars and trucks was taken from Automotive News Market Data Books for 1979–1985. The sample consists of the base versions (i.e., without options) of every model available each year, with the exception of station wagons. In addition to the suggested retail price in March or April of each year, data were obtained on the quantity sold and various characteristics. In Table I we report the number of models, unit-values, and price indexes for the sample. For comparison, the U. S. consumer price index is shown in the last row.

Considering the data for cars, the rise in unit-value over 1980–1981 is measured as 20 percent in nominal terms, or about 10 percent in real terms. This movement reflects both price increases for individual models (which may be upgraded), and any shift in demand toward higher priced cars. In the fourth row of Table I we report a price index for cars constructed using constant index-weights between each pair of years. Movements in this index reflect price rises only for individual models.

Turning to compact trucks, the retail prices in Table I include the effects of the higher ad valorem tariff beginning in August 1980. We observe a 28 percent nominal increase in the unit-value or price index over 1980–1981. For later purposes, it will be useful to estimate what amount of this price rise is due to the tariff. The rate of duty was increased from 4 to 25 percent, applied to the wholesale (f.a.s.) prices. The wholesale values for trucks are about three

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9. This unusually high rate was a result of the "chicken war" between the United States and Europe in 1962–1963, when the United States retaliated against higher tariffs on poultry sales to West Germany. See USITC [1980, pp. A10–A11].

10. The index numbers reported are the Fisher Ideal form, which is the geometric mean of the Laspeyres and Paasche indexes. The exact construction of these indexes, including a discussion of how new models are handled, is in Feenstra [1985, sec. 6].
# Table I
## Sample of Japanese Autos

<table>
<thead>
<tr>
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</tr>
</thead>
<tbody>
<tr>
<td><strong>Cars</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Number of models</td>
<td>21</td>
<td>24</td>
<td>24</td>
<td>24</td>
<td>26</td>
<td>29</td>
<td>31</td>
</tr>
<tr>
<td>Unit-value (percent change)</td>
<td>4,946</td>
<td>5,175</td>
<td>6,211</td>
<td>6,834</td>
<td>7,069</td>
<td>7,518</td>
<td>8,038</td>
</tr>
<tr>
<td>Price index (percent change)</td>
<td>96.8</td>
<td>100.0</td>
<td>119.8</td>
<td>129.1</td>
<td>131.2</td>
<td>138.8</td>
<td>148.3</td>
</tr>
<tr>
<td>Unit-quality (percent change)</td>
<td>5,002</td>
<td>5,124</td>
<td>5,511</td>
<td>5,896</td>
<td>6,257</td>
<td>6,488</td>
<td>6,709</td>
</tr>
<tr>
<td>Quality index (percent change)</td>
<td>98.7</td>
<td>100.0</td>
<td>107.4</td>
<td>112.8</td>
<td>117.3</td>
<td>121.3</td>
<td>125.4</td>
</tr>
<tr>
<td><strong>Trucks</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Number of models</td>
<td>10</td>
<td>10</td>
<td>11</td>
<td>11</td>
<td>10</td>
<td>11</td>
<td>12</td>
</tr>
<tr>
<td>Unit-value (percent change)</td>
<td>4,804</td>
<td>4,937</td>
<td>6,298</td>
<td>6,419</td>
<td>6,089</td>
<td>6,261</td>
<td>6,339</td>
</tr>
<tr>
<td>Price index (percent change)</td>
<td>97.2</td>
<td>100.0</td>
<td>127.5</td>
<td>131.8</td>
<td>125.3</td>
<td>127.3</td>
<td>129.7</td>
</tr>
<tr>
<td>Unit-quality (percent change)</td>
<td>4,952</td>
<td>4,943</td>
<td>5,041</td>
<td>5,275</td>
<td>5,276</td>
<td>5,274</td>
<td>5,620</td>
</tr>
<tr>
<td>Quality index (percent change)</td>
<td>99.8</td>
<td>100.0</td>
<td>102.1</td>
<td>106.4</td>
<td>105.5</td>
<td>104.7</td>
<td>112.3</td>
</tr>
<tr>
<td>U.S. consumer prices (percent change)</td>
<td>88.1</td>
<td>100.0</td>
<td>110.4</td>
<td>117.1</td>
<td>120.9</td>
<td>126.1</td>
<td>130.6</td>
</tr>
</tbody>
</table>

*Includes utility vehicles.*
quarters of the suggested retail values (net of duty). Thus, a rise of 21 percent in the tariff at the wholesale level corresponds to a 16 percent price rise at the retail level. We shall use this figure below.

IV. Hedonic Regressions

Our objective in the next sections is to measure the quality of Japanese auto imports using hedonic regressions [Griliches, 1971]. The hedonic regression is an estimate of the equilibrium price schedule \( p(z) \). In this estimation we shall pool data over 1979–1985. Thus, the hedonic regressions should include coefficients allowing prices to rise proportionately over time (due to changes in costs or exchange rates), and by specific dollar amounts over time (due to the quota in cars). Our discussion in Section II suggested that in the short run, the quota could also have nonlinear effects on the equilibrium price schedule \( p(z) \). However, in other estimation we have formally accepted the hypothesis that, aside from proportional and specific price changes, the hedonic regressions for Japanese cars and trucks are stable over time [Feenstra, 1985, 1988]. This result is consistent with the long-run equilibrium price schedule in (6), so we can interpret our regressions as estimating this long-run schedule.

We specify the hedonic regressions for cars and trucks as follows:

\[
\begin{align*}
  p_{tk} &= s_t + \exp(\delta_t + \gamma'z_{tk}) + \epsilon_{tk}^e \\
  p_{tk}^* &= \exp(\delta_t^* + \gamma'^*z_{tk}^*) + \epsilon_{tk}^*
\end{align*}
\]

where \( t = 79, \ldots, 85 \) denotes years, \( k \) denotes models, (10a) refers to Japanese cars, and (10b) to Japanese trucks. The functional form in (10) is common in empirical applications (see Atkinson and Halvorsen [1984], who test for alternatives). The coefficients \( s_t \) are the specific price effect of the quota in cars, which are assumed to be equal across firms as in (6). Proportional movement in prices, as well as the ad valorem tariff in trucks, is reflected in \( \delta_t \) and \( \delta_t^* \).

The OLS residuals from (10) showed very significant evidence

11. Using data from Feenstra [1986], the ratios of wholesale to retail values (net of duty) lie in the range 0.742–0.789 for 1979–1985.

12. This result does not hold for American cars, however, since Feenstra [1985] finds evidence of a structural change over 1980–1981 in the American small car regression. A similar conclusion is obtained on a more comprehensive data set by Ohta and Griliches [1983], who attribute the structural change to the rise in gasoline prices.
of heteroskedasticity. To correct for this, we shall use generalized least squares, as follows: the regressions are first estimated with OLS to obtain a measure of \( \exp(\hat{\delta}_t + \hat{\gamma}'z_{tk}) \) and similarly for trucks, and then reestimated while weighting each observation by the inverse of \( \exp(\hat{\delta}_t + \hat{\gamma}'z_{tk}) \). As discussed by Eppl [1987], the errors in (10) can arise from missing or proxied characteristics, so \( z_{tk} \) may be correlated with \( \epsilon_{tk} \). By using least squares, the estimates \( \hat{\gamma} \) and \( \hat{\gamma}^* \) will reflect both the true coefficient of each characteristic and its correlation with missing variables. However, since our reason for estimating (10) is to obtain a measure of product quality, including if possible the contribution of missing characteristics, then the bias in \( \hat{\gamma} \) and \( \hat{\gamma}^* \) seems to be acceptable.

Product quality is defined as \( \exp(\delta_{79} + \gamma'z_{tk}) \) and similarly for trucks, i.e., the predicted price from the hedonic regression not including the contribution of year dummies. Given this measure of product quality, we should check that it is not too sensitive to the list of characteristics used. We shall follow the approach of Leamer [1983], and estimate (10) for various lists of characteristics. In Section V below, we report the ranges of product quality obtained in this manner.

The initial estimates of (10), using generalized least squares, are presented in the first two columns of Table II. Each estimated coefficient can be interpreted as the percentage change in price caused by a unit change in that characteristic; e.g., adding one foot of width to a car leads to a 35 percent rise in price. The presence of some negative coefficients is not surprising given the multicollinearity in our data. The estimates \( \hat{\delta}_t \) and \( \hat{\delta}^*_t \) are shown as the “proportional” year coefficients in the lower half of Table II, while \( \hat{s}_t \) are shown as the “specific” year coefficients. In the bottom of column 1, \( \hat{s}_t \) are positive for 1980–1984 and of reasonable magnitude, but totally insignificant. This occurs because of the multicollinearity between the specific and proportional year dummies in (10a).

To improve the estimates of \( s_t \), we consider pooling the car and truck data allowing for different characteristic coefficients, while

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13. Ranking cars by their estimated quality, we selected sixteen large models, fourteen small models, and omitted five intermediate cases. The OLS estimates of (10a) for large and small models gave standard errors of $746 and $426, respectively. The \( F \)-statistic for testing the null hypothesis of equal error variances is computed as 2.97, which compares with \( F_{0.99}(62,60) = 1.84 \), so we have strong evidence of heteroskedasticity.
### TABLE II
HEDONIC REGRESSIONS, DEPENDENT VARIABLE—PRICE

<table>
<thead>
<tr>
<th>Sample</th>
<th>Cars</th>
<th>Trucks</th>
<th>Cars</th>
<th>Trucks</th>
</tr>
</thead>
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<tr>
<td>Observations</td>
<td>179</td>
<td>75</td>
<td>254</td>
<td>0.917</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.900</td>
<td>0.956</td>
<td>0.102</td>
<td>0.101</td>
</tr>
<tr>
<td>Standard error</td>
<td>0.102</td>
<td>0.077</td>
<td>0.101</td>
<td>0.101</td>
</tr>
<tr>
<td>Constant</td>
<td>6.35*</td>
<td>7.18*</td>
<td>6.80*</td>
<td>7.01*</td>
</tr>
<tr>
<td>(0.58)</td>
<td>(0.57)</td>
<td>(0.53)</td>
<td>(0.73)</td>
<td></td>
</tr>
<tr>
<td>Weight (tons)</td>
<td>0.034</td>
<td>0.30*</td>
<td>0.002</td>
<td>0.32*</td>
</tr>
<tr>
<td>(0.120)</td>
<td>(0.080)</td>
<td>(0.108)</td>
<td>(0.104)</td>
<td></td>
</tr>
<tr>
<td>Width (feet)</td>
<td>0.35*</td>
<td>0.058</td>
<td>0.35*</td>
<td>0.11</td>
</tr>
<tr>
<td>(0.11)</td>
<td>(0.115)</td>
<td>(0.097)</td>
<td>(0.14)</td>
<td></td>
</tr>
<tr>
<td>Height (feet)</td>
<td>−0.064</td>
<td>0.12*</td>
<td>−0.16*</td>
<td>0.991</td>
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<td>(0.055)</td>
<td>(0.039)</td>
<td>(0.067)</td>
<td>(0.050)</td>
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<tr>
<td>Horsepower (100)</td>
<td>0.69*</td>
<td>0.010</td>
<td>0.70*</td>
<td>0.066</td>
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<tr>
<td>(0.085)</td>
<td>(0.109)</td>
<td>(0.072)</td>
<td>(0.142)</td>
<td></td>
</tr>
<tr>
<td>Transmission</td>
<td>0.14*</td>
<td>0.046</td>
<td>0.16*</td>
<td>0.081*</td>
</tr>
<tr>
<td>(5-speed or auto)</td>
<td>(0.024)</td>
<td>(0.026)</td>
<td>(0.021)</td>
<td>(0.031)</td>
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<tr>
<td>Power steering</td>
<td>0.058*</td>
<td>0.041</td>
<td>0.073*</td>
<td>0.008</td>
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<tr>
<td>(0.027)</td>
<td>(0.046)</td>
<td>(0.025)</td>
<td>(0.054)</td>
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</tr>
<tr>
<td>Air conditioning</td>
<td>0.15*</td>
<td>0.17*</td>
<td>0.17*</td>
<td>0.031</td>
</tr>
<tr>
<td>(0.033)</td>
<td>(0.033)</td>
<td>(0.031)</td>
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<tr>
<td>Four-wheel drive</td>
<td></td>
<td>0.17*</td>
<td>0.17*</td>
<td>0.21*</td>
</tr>
<tr>
<td>(0.031)</td>
<td>(0.031)</td>
<td>(0.038)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Proportional</td>
<td></td>
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<td></td>
</tr>
<tr>
<td>Year 1980</td>
<td>0.009</td>
<td>0.024</td>
<td>0.011</td>
<td>0.011</td>
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<tr>
<td>(0.031)</td>
<td>(0.034)</td>
<td>(0.025)</td>
<td>(0.025)</td>
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</tr>
<tr>
<td>Year 1981</td>
<td>0.047</td>
<td>0.225*</td>
<td>0.049</td>
<td>0.209*</td>
</tr>
<tr>
<td>(0.109)</td>
<td>(0.033)</td>
<td>(0.040)</td>
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<tr>
<td>Year 1982</td>
<td>0.095</td>
<td>0.236*</td>
<td>0.045</td>
<td>0.205*</td>
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<tr>
<td>(0.108)</td>
<td>(0.035)</td>
<td>(0.041)</td>
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<tr>
<td>Year 1983</td>
<td>0.021</td>
<td>0.184*</td>
<td>−0.024</td>
<td>0.136*</td>
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<tr>
<td>(0.105)</td>
<td>(0.037)</td>
<td>(0.044)</td>
<td>(0.044)</td>
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<tr>
<td>Year 1984</td>
<td>0.094</td>
<td>0.218*</td>
<td>0.016</td>
<td>0.176*</td>
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<tr>
<td>(0.102)</td>
<td>(0.037)</td>
<td>(0.044)</td>
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<tr>
<td>Year 1985</td>
<td>0.221*</td>
<td>0.217*</td>
<td>0.169*</td>
<td>0.100</td>
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<tr>
<td>(0.101)</td>
<td>(0.037)</td>
<td>(0.069)</td>
<td>(0.051)</td>
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</tr>
<tr>
<td>Specific</td>
<td></td>
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<td></td>
</tr>
<tr>
<td>Year 1981</td>
<td>414</td>
<td>434</td>
<td>434</td>
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</tr>
<tr>
<td>(618)</td>
<td>(250)</td>
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<tr>
<td>Year 1982</td>
<td>376</td>
<td>707*</td>
<td>707*</td>
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</tr>
<tr>
<td>(650)</td>
<td>(256)</td>
<td></td>
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<tr>
<td>Year 1983</td>
<td>768</td>
<td>1,085*</td>
<td>1,085*</td>
<td></td>
</tr>
<tr>
<td>(601)</td>
<td>(262)</td>
<td></td>
<td></td>
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<tr>
<td>Year 1984</td>
<td>617</td>
<td>1,096*</td>
<td>1,096*</td>
<td></td>
</tr>
<tr>
<td>(620)</td>
<td>(267)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Year 1985</td>
<td>−131</td>
<td>256</td>
<td>256</td>
<td></td>
</tr>
<tr>
<td>(720)</td>
<td>(492)</td>
<td></td>
<td></td>
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</tr>
</tbody>
</table>

*Significant at 95 percent level. Standard errors are in parentheses.
testing the following cross-equation constraints:

\[
\delta_{80} = \delta^*_80, \\
\delta_t = \delta^*_t - 0.16, \quad t = 81, 82, \ldots, 85.
\]

We interpret \(\delta_t\) as measuring the "quality-adjusted" price change that would have occurred in cars without the trade restraint. Then (11) says that the quota-free price change in cars, from 1981 onward, would have been 16 percent less than that in trucks, where this 16 percent is the increase in the ad valorem tariff on trucks at a retail level. We are thus testing whether the yearly change in truck prices, corrected for quality and the tariff, equals the yearly change in car prices, corrected for quality and the quota.\textsuperscript{14} This approach is based on the idea that cars and trucks are faced with similar cost conditions in Japan, and identical exchange rates when converting to dollars.

When the constraints (11) are tested, they can be accepted for 1980–1984 but not for 1985.\textsuperscript{15} This result can be seen by comparing the point estimates of \(\delta_t\) and \(\delta^*_t\) in columns 1 and 2 of Table II, which differ by roughly 0.16 in 1981–1984 but are nearly equal in 1985. Apparently, Japanese cars experienced a proportional price rise in 1984–1985, which was not reflected in truck prices. This result is confirmed by examining individual models, where a number of luxury Japanese cars had quality adjusted, percentage price increases over 1984–1985 which were equal to or exceeded those of economy models. A possible explanation for this finding is discussed in the next section.

The car and truck regressions with (11) imposed for 1980–1984 are reported in the last two columns of Table II. Of principal interest are the specific price changes \(\hat{s}\) for cars, in the second to last column. For 1982–1984 these estimates are large, rising, and highly significant. The technique of pooling the car and truck data is very effective in reducing the standard errors of \(\hat{s}\). These specific price changes are the estimated effect of the quota restraint in cars, which exceeds $1,000 in 1983 and 1984 with a standard error of $270.

\textsuperscript{14} For utility vehicles subject to the quota, we use the characteristic coefficients of the truck regression and the year coefficients of the car regression.

\textsuperscript{15} Using OLS, the SSR from (10) is 89.03 (million) without (11), and 93.00 with (11) imposed for 1980–1984. The \(F\)-statistic is computed as 1.97 which compares with \(F_{0.95}(5, 221) = 2.25\), so we can accept the constraints (11) over 1980–1984. However, when they are imposed for 1980–1985, we obtained an SSR of 95.12, and the \(F\)-statistic is 2.52. This compares with \(F_{0.95}(6, 221) = 2.14\), leading to a rejection of (11) over the 1980–1985 period. This rejection is also obtained if we test the 1985 constraint conditional on the 1980–1984 constraints being imposed.
V. RESULTS AND CONCLUSIONS

A. Quality Estimates

Our estimates of quality change for Japanese cars and trucks are reported in Table I. The quality of each model is measured as the predicted price from the hedonic regression (columns 3 and 4 in Table II) not including the contribution of the year dummies. The unit-quality is then computed as a weighted average across models, using the quantities sold in that year as weights. Movements in the unit-quality reflect both quality change in individual models and shifts in demand across models. It can be interpreted as the change in average quality discussed in subsection IIA. The quality index is calculated by aggregating the quality of individual models, using constant weights between each pair of years.\textsuperscript{16} It can be interpreted as measuring the (short-run) response of firms, i.e., as the change in $p_{t,z}$ due to the quota or tariff.

For Japanese cars in Table I we find a 31 percent rise in unit-quality and a 25.4 percent rise in the quality index over 1980–1985. This is about one half of the nominal increase in car prices. The quality upgrading is most evident in 1981, the first year of the trade restraint.\textsuperscript{17} It is interesting that the movements in unit-quality over time are similar to the quality index, with the unit-quality rising somewhat faster. This suggests that consumers not only accepted the quality upgrading of firms, but actually moved up in their choice of models. Feenstra [1984] identified three 1981 models that were the highest priced, had the greatest quality upgrading, and also experienced a substantial increase in sales over 1980–1982: the Toyota Cressida, Toyota Supra, and Datson 810 Maxima. These models also had reductions in their quality-adjusted prices, which explains the response of consumers.

In the nomenclature of Leamer [1983], the precise list of characteristics in our hedonic regression is “doubtful.” We checked that our quality measures are not too sensitive to this list by reestimating (10), omitting each of the following variables: length, weight, width, height, and horsepower.\textsuperscript{18} The following ranges of unit-quality and the quality index for cars are obtained:

\textsuperscript{16} See footnote 10.

\textsuperscript{17} The price and characteristics data for 1981 were collected on April 10, before the formal announcement of the voluntary export restraint on May 1. In attributing the 1980–1981 quality upgrading to the trade restraint, we are supposing that Japanese producers anticipated the quota.

\textsuperscript{18} Length is already omitted in Table II, since it is highly insignificant when used. The constraints (11) were accepted and imposed for 1980–1984 in each of the other regressions in the sensitivity exercise.
The lower (upper) end of the quality ranges are obtained when horsepower (height) is omitted. While some variation in the quality estimates is obtained, the hypothesis of quality upgrading in Japanese cars is still confirmed.

Turning to Japanese trucks, in the bottom of Table I we report a 4.7 percent rise in the quality index over 1980–1984, with a further 7.3 percent rise in 1984–1985. Between 1980 and 1984 the quality index increases in two years and falls slightly in two others. The rise of 4.2 percent over 1981–1982 can be primarily attributed to the addition of five-speed transmissions on several models. The quality upgrading of 7.3 percent over 1984–1985 reflects a recent movement toward heavier specification of Japanese trucks. Overall, there is little evidence of any sustained quality change due to the rise in the ad valorem tariff. These results for cars and trucks are in substantial agreement with the theory in Section II and existing literature.

B. Price Effect of the Quota

By pooling our car and truck data, we have obtained accurate estimates of the specific price increases in cars due to the quota (Table II, bottom of column 3). The ranges of specific price rises obtained from our sensitivity exercise are reported in the tabulation above, and exceed $1,000 for 1983 and 1984. These price increases can be multiplied by the quantity of cars imported to obtain a social cost of about $2 billion annually in recent years. This is the rectangle of quota rents transferred to Japanese retailers and producers. The additional deadweight loss triangle would be a fraction of the quota rents. For example, the specific quota increase accounts for about 15 percent of the price of cars in 1983–1984, which would reduce sales by some 600,000 units using an elasticity of −2. The deadweight loss would then be $300 million, leading to a total welfare cost of $2.3 billion annually due to the quota.  

19. Our welfare results are similar to those in Hufbauer, Berliner, and Elliot [1986], who survey other studies. Crandall [1984] estimates the effect of the quota on U.S. domestic car prices, while Dinopoulos and Kreinin [1987] consider the effect on European car prices.
Finally, we are left with the puzzle of why Japanese car prices experienced a large proportional price increase over 1984–1985, unlike the behavior of truck prices. An intriguing possibility is that the quota has facilitated more collusive behavior by Japanese firms, as proposed in theory by Krishna [1984] and Harris [1985]. Dixit [1988] finds some evidence of more collusive behavior by Japanese firms in the early years of the quota. Bresnahan [1981] has demonstrated that monopolistic pricing in the U.S. auto industry leads to higher percentage markups on more expensive cars, and this is the direction toward which the Japanese prices have changed recently. If this pattern persists, it suggests some long-run shift in the pricing strategy of Japanese firms, reinforced by the continued use of the trade restraint by the Japanese government. Together with the clear shift in product quality, these results argue that the effects of the quota will be difficult to reverse in future years.

REFERENCES


20. Another possibility is that the reallocation of the quota in 1985, which gave smaller percentage increases to Toyota and Nissan than to other producers, resulted in specific price changes that differed across firms. Our analysis of residuals from the hedonic regressions did not show evidence of this, however.

—, and —, "Oligopolistic Competition and International Trade: Quantity and Quality Restrictions," mimeo, University of Wisconsin—Milwaukee and Tel Aviv University, 1986.


