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환율전이와 시장의 반응: 미국 철강시장에서의 한국과 일본의 경쟁

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Exchange Rate Pass-Through and Market Response: Competition between Korea and Japan in the US Steel Market

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ABSTRACT

This paper theoretically formulated and empirically explored the relationship between exchange rate pass-through (ERPT) for (average) market price and an individual country's price, using steel products data in the US market, with special reference to two major steel exporting countries, Korea and Japan. It was found that the direction of market ERPT can be different from that of individual ERPT that each exporter experiences, due to strategic interactions among producers and different parameters. Vector error correction (VEC) models and impulse response analysis were used with the statistical inference based on the bootstrap-after- bootstrap of Kilian (1998) for short-run, and the fully modified estimation of Phillips and Hansen (1990) was used for long-run. Empirical results indicate that market ERPT in the US market due to changes in Korea-US exchange rates is different from those due to changes in Japan-US exchange rates. The framework developed in this study indicates that this phenomenon is attributed to either (i) the two countries have individual ERPTs of different magnitudes and directions for the products in the US market, or (ii) the pricing strategies of the other exporters' (to the US steel market) respond differently depending on whether the price of the product from Korea changes or that from Japan does. As each exporter's ERPT can be significantly different, and market response to each country's ERPT can be also different, this study concludes that it is crucial for an exporter to understand how competitors in the market respond to changes in its price, as well as to understand how its price changes when the relevant exchange rate fluctuates.

환율의 변화에 따른 개별 교역재 가격의 변화는 ‘환율전이효과(Exchange rate pass-through)’로 불리며, 국제경제학의 가장 중요한 연구분야 중의 하나로 인식되고 있다. 또한 환율변화 당사국의 교역재 가격변화에 따라 시장전체에서 일어나는 환율의 전이현상은, 수출시장과 특정수출국의 환율이 변화할 때 수출국의 가격책정전략과 시장전체의 반응에 대한 정보를 제공해줄 수 있다는 점에서 깊은 연구가 필요하다. 그러나 대부분의 연구는 환율변화에 따른 특정 국가의 수출재 가격변화만을 고려할 뿐, 경쟁국 재화가격의 변화나, 이에 따른 수출대상국의 재화시장이 전체적으로 받는 충격에 대한 연구는 거의 이루어지지 않고 있다. 본 논문은, 한국과 일본의 대미 달러 환율변화에 따른 철강재의 가격변화를 통해 시장의 반응을 분석한다. Bootstrap-after-bootstrap을 활용한 vector error correction 모형과 이에 따른 충격반응함수분석, 그리고 Phillips-Hansen의 추정방법(fully modified estimation)을 통한 분석은 몇 가지 중요한 시사점을 제공하는데, 그중 가장 중요한 두 가지를 요약하면 다음과 같다. 즉, (i) 한일 양국의 환율변화는 미국 철강시장에서의 양국 수출철강가격에 대해 서로 다른 개별적 전이효과를 야기하며, (ii) 여타 수출국의 반응도 한일 양국 중 어느 나라 제품의 가격이 변화하는가에 따라 달라진다는 것이다. 이에 따라 같은 비율의 환율변화에 대해서도 그 당사국에 따라 시장의 가격반응은 달라지게 됨을 알 수 있다.

I. Introduction

The post-Bretton Woods era that allowed the free fluctuation of exchange rates provided the impetus for research on the effect of exchange rate shocks on commodity prices. This topic was explored more intensively in the 1980s as economies experienced an unprecedented fluctuation in real exchange rates accompanied by the appreciation (until 1985), and the subsequent depreciation of the US dollar in the same decade. While the first few years of the 1990s have been characterized as a period of stability in foreign exchange markets as Goldberg and Knetter (1997) point out, there were still some notably large fluctuations in various currency values. In the 1980's and 1990's, it was frequently observed that the price of commodities in an importing country did not fluctuate as expected or predicted by the traditional models such as the law of one price.

The incomplete pass-through (when changes in the exchange rates are not fully transferred to commodity prices), or perverse pass-through (when changes in the exchange rate influence commodity prices in unexpected ways), observed in these periods was, in general, attributed by researchers to the fact that foreign producers may respond to a dollar appreciation by partially decreasing their prices and also increasing their profit margins. On the other hand, in periods of dollar depreciation it was observed that they may increase their prices but also reduce their profit margins, in order to keep up sales and defend their market share (for example, Gagnon & Knetter, 1995; Krugman, 1987; Tivig, 1996; Varangis & Duncan, 1993).

Notwithstanding some unexpected outcomes such as perverse movement or no pass-through of commodity prices, most studies that utilized disaggregated data (such as 4-digit country specific industry data) reported the existence of pass-through. However, the extent of pass-through was partial and differentiated by periods and market structure, across regions and products (for example, Feenstra, 1989; Feenstra, Gagnon & Knetter, 1996; Gagnon & Knetter, 1995; Knetter, 1989, 1995; Marston, 1990).

The purpose of this paper is to explore the relationship between exchange rate pass-through (hereafter ERPT) for individual producers and for the market. While studies on ERPT for individual producers have been extensively carried out, it is surprising that the effect of changes in a specific exchange rate on the market price (hereafter *market exchange rate pass-through* or *market ERPT*) has seldom been examined.[1] A few exceptions include Rindler and Yandle (1972), Sjaastad

[1] In fact, when the law of one price prevails or the market is perfectly competitive, the mar-

(1985), Sjaastad and Scacciavilani (1996), and Tcha and Sjaastad (1998). While these studies explored the effect of one-period shocks on commodity markets in a multi-period open economy, they did not examine the mechanism of how the individual ERPT is related to the market ERPT. In imperfectly competitive markets, all competitors' prices (and therefore market shares) are closely related to one another by strategic interaction, hence ERPT for the whole market can be different from that for an individual producer experiencing exchange rate fluctuations.

Due to the nature of oligopoly markets or implicit cartels, it is sometimes difficult to find each individual producer's real price response to a given change in a certain exchange rate, especially where only the weighted average market price is announced. Different categorization of products across countries also hinders the direct observation of individual prices. Moreover, the analysis of the response of the market price to exchange rate shocks will provide useful market information that the analysis of individual producer's ERPT cannot provide. This paper theoretically develops a model, which explains the relationship between the individual ERPT and the market ERPT, by accommodating the aforementioned problems associated with market prices and data deficiencies. The theoretical framework to analyze individual producers behavior is based on Froot and Klemperer (1989) and Tivig (1996), which were widely applied in studies on ERPT for individual producers, such as Gross and Schmitt (2000). This study goes one step further, by analyzing the market ERPT and its relationship with the individual ERPT. Necessary and sufficient conditions to define the direction of movements of market and individual prices will be explored. Further discussions on this subject will be introduced after ERPT for the US steel market is empirically analyzed.

II. The Market Exchange Rate Pass-Through

1. Market ERPT and Individual ERPT

The weighted average market price of a steel product that we are interested in is $P_t = \sum_{i=1}^{n} \omega_{it} P_{it}$ at time t, where ω_i is country i's mar-

ket ERPT will provide the same results as the individual ERPT. However, when the market is imperfect, the change of average price in the market is not necessarily identical to that of the price of a good from the country whose exchange rate varies.

ket share and P_i is country i's price in the US market, which is determined from profit maximization process of exporter i. The effect of changes in the exchange rate between country J and the US (e_J) on the weighted average market price is, leaving out time subscript t for simplicity,

$$\frac{\partial P}{\partial e_J} = \frac{\partial P}{\partial P_J}\frac{\partial P_J}{\partial e_J} = \sum_{i=1}^{n}\left[\frac{\partial \omega_i}{\partial P_i}\frac{\partial P_i}{\partial P_J} P_i + \frac{\partial P_i}{\partial P_J}\omega_i\right]\frac{\partial P_J}{\partial e_J}. \quad (1)$$

Due to imperfect competition in the market, both P_i and ω_i are affected by P_J, unless each country's market share equi-proportionally changes with total demand when P_J changes. Therefore, using elasticities, equation (1) can be rewritten as

$$\begin{aligned}\frac{\partial P}{\partial e_J} &= \frac{\partial P_J}{\partial e_J}\sum_{i=1}^{n}\left(\frac{\partial \omega_i}{\partial P_i}\frac{P_i}{\omega_i}\frac{\partial P_i}{\partial P_J}\frac{P_J}{P_i}\cdot\frac{\omega_i P_i}{P_J} + \frac{\partial P_i}{\partial P_J}\frac{P_J}{P_i}\frac{\omega_i P_i}{P_J}\right)\\ &= \frac{\partial P_J}{\partial e_J}\omega_J\left(\sum_{i=1}^{n}(1-\phi_{ii})\,\sigma_{iJ}\cdot\frac{\omega_i P_i}{\omega_J P_J}\right), \qquad (2)\end{aligned}$$

where ϕ_{ii} is the elasticity of country i's market share with respect to P_i $\left(i.e.,\ \phi_{ii} = -\frac{\partial \omega_i}{\partial P_i}\frac{P_i}{\omega_i}\right)$, and σ_{iJ} is the cross elasticity of country i's price P_i with respect to P_J $\left(i.e.,\ \sigma_{iJ} = \frac{\partial P_i}{\partial P_J}\frac{P_J}{P_i}\right)$.[2]

Separating country J's own elasticities (ϕ_{JJ} and σ_{JJ}, where $\sigma_{JJ}=1$) from (2) gives

$$\frac{\partial P}{\partial e_J} = \frac{\partial P_J}{\partial e_J}\omega_J\left[1-\phi_{JJ} + \sum_{i\neq J}(1-\phi_{ii})\,\sigma_{iJ}\frac{\omega_i P_i}{\omega_J P_J}\right]. \qquad (3)$$

Equations (1) and (3) show that $[\partial P/\partial P_J]$ is the important chain connecting the market ERPT and country J's ERPT when e_J changes. This term $[\partial P/\partial P_J]$ is, in turn, the product of country J's market share

[2] We do not have any presumption about the sign of this cross elasticity. Country i may increases its price when country J increases its price. Alternatively, it may be the case that country i decreases its price as country J increases its price. It depends on country i's profit maximization behavior.

and the constellation of parameters and variables in the bracket as shown in (3). Since country J's market share ω_J is positive, the sign of $[\partial P/\partial e_J]$ in (3) is determined by the sign of the bracket on the right side such as

$$sign\left[\frac{\partial P}{\partial e_J}\right] = \begin{cases} sign\left[\dfrac{\partial P_J}{\partial e_J}\right] & iff\ \left[1 - \phi_{JJ} + \sum_{i \neq J}(1 - \phi_{ii})\ \sigma_{iJ}\ \dfrac{\omega_i P_i}{\omega_J P_J}\right] > 0 \\ -sign\left[\dfrac{\partial P_J}{\partial e_J}\right] & iff\ \left[1 - \phi_{JJ} + \sum_{i \neq J}(1 - \phi_{ii})\ \sigma_{iJ}\ \dfrac{\omega_i P_i}{\omega_J P_J}\right] < 0 \end{cases} \tag{4}$$

Equation (4) indicates that the market price does not necessarily move in the same direction as country J's price when the exchange rate between country J and the destination (e_J) changes temporarily. An assumption helps to explore the necessary and sufficient conditions in (4) further.

[Assumption 1] The total market demand for a specific steel product (M) is not affected by the change in a producer's price (i.e., $\mu_J = \dfrac{\partial M}{\partial P_J}\dfrac{P_J}{M} = 0$, *where* μ_J *is the elasticity of market demand with respect to J's price).* [3]

Applying this assumption to the equation for the elasticity of the market share for country J, ϕ_{JJ}, provides

$$\phi_{JJ} = \frac{\partial\ (M_J/M)}{\partial\ P_J}\ \frac{P_J}{(M_J/M)} = \frac{\partial M_J}{\partial P_J}\ \frac{P_J}{M_J} - \frac{\partial M}{\partial P_J}\ \frac{P_J}{M} = \varepsilon_{JJ} - \mu_J = \varepsilon_{JJ},$$

where ε_{JJ} is the elasticity of demand for country J $\Big[i.e.\ \varepsilon_{JJ} =$

[3] While this assumption is consistent with a Cobb-Douglas function with powers as weights (shares) of each exporter and, in general, reasonable for the factor markets investigated in this paper, the total market demand for final goods would change as prices change. This assumption is needed for the convenience of the analysis only, and does not change the findings in this paper significantly.

$\left.\frac{\partial M_J}{\partial P_J}\frac{P_J}{M_J}\right]$. Therefore, the term that determines the sign of $[\partial P/\partial P_J]$ is from (4) converted to

$$1 - \phi_{JJ} + \sum_{i \neq J}(1 - \phi_{ii})\ \sigma_{iJ}\ \frac{\omega_i P_i}{\omega_J P_J} =$$

$$1 - \varepsilon_{JJ} + \sum_{i \neq J}(1 - \varepsilon_{ii})\ \sigma_{iJ}\ \frac{\omega_i P_i}{\omega_J P_J}. \tag{5}$$

Assumption 1 also implies that country J's market loss due to its price change is all absorbed by the other countries, i.e. $\left[\frac{\partial M_J}{\partial P_J} = -\sum_{i \neq J}\frac{\partial M_i}{\partial P_J}\right]$, hence the elasticity of demand for country J's products is

$$\varepsilon_{JJ} = \frac{\partial M_J}{\partial P_J}\ \frac{P_J}{M_J} = -\sum_{i \neq J}\frac{\partial M_i}{\partial P_J}\ \frac{P_J}{M_J} = -\sum_{i \neq J}\varepsilon_{ii}\ \sigma_{iJ}\ \frac{M_i}{M_J}. \tag{6}$$

Substituting (6) for ε_{JJ} in (5) gives

$$1 - \varepsilon_{JJ} + \sum_{i \neq J}(1 - \varepsilon_{ii})\ \sigma_{iJ}\ \frac{\omega_i P_i}{\omega_J P_J} =$$

$$1 + \sum_{i \neq J}\left[\ \sigma_{iJ}\ \left(1 - \varepsilon_{ii} + \varepsilon_{ii}\frac{P_J}{P_i}\right)\right]\ \frac{\omega_i P_i}{\omega_J P_J}, \tag{5)'$$

which indicates that, when there are n countries in the market, the direction of the market ERPT to the change in any country's exchange rate depends on the magnitude and signs of all the other countries' elasticities, prices and market shares as well as country J's ERPT.

We will first examine an extreme and simple case where all n suppliers are identical, in the sense that their elasticities, prices and market shares are initially symmetric, and then turn to look at a general case, where firms are allowed to be different. It will be shown that while both necessary and sufficient conditions are obtained in a symmetric case, only sufficient conditions can be derived in a more general case.

2. Market ERPT with Identical Producers

When all firms are identical and share the same variables and parameters, condition (5)' reduces to $[1 + (n - 1)\ \sigma]$ and, accordingly, the ERPT for the market is

$$\frac{\partial P}{\partial e_J} = \frac{\partial P_J}{\partial e_J}\ \omega_J\ \left[1 + (n - 1)\ \sigma\right].$$

The following proposition is presented regarding the necessary and sufficient conditions for the direction of the price movements.

[Proposition 1] When all the producers in the market are identical, $\sigma \underset{(<)}{>} -\dfrac{1}{(n - 1)}$ *is the necessary and sufficient condition for country J's and the market prices to move to the same (opposite) direction(s), or the individual and the market ERPT have the same (different) signs, when the exchange rate between country J and the market changes*[4].

The implication of this condition is clear. It shows that even when the prices of all the other producers move in the opposite direction to country *J*'s price due to their dynamic profit maximization strategy, it is still possible that the direction of the market ERPT is the same as that of country *J*'s, depending on the elasticity of country *i*'s price with respect to country *J*'s. In the duopolist case, for example, only if elasticity is greater than -1 will the market price move in the same direction as country *J*'s, regardless of whether country *J* responds normally or perversely, and whether all of the other countries respond normally and perversely. As the number of suppliers increases, $[-1/(n-1)]$ approaches zero, that implies that the two ERPTs (say, country *J*'s and the US market's) are less likely to have the same sign unless the cross elasticity is positive.

3. Market ERPT with Non-identical Producers

When producers are not identical, it is not possible to derive the sufficient and necessary conditions for the direction of market ERPT and country *J*'s ERPT in a more succinct form. Nonetheless, (5)' is always positive if the bracket is always bigger than zero for all i. As P_i,

[4] It is open to question whether other countries would move in a different direction from country *J*. Tivig (1996) theoretically proves that σ_{iJ} is always positive if producers operate in the inelastic region, while the sign of σ_{iJ} is indeterminate in the elastic region.

M_i , P_J and M_J are all positive, the sufficient conditions to have $sign\left[\frac{\partial P}{\partial e_J}\right] = sign\left[\frac{\partial P_J}{\partial e_J}\right]$ or $sign\left[\frac{\partial P}{\partial e_J}\right] = -\,sign\left[\frac{\partial P_J}{\partial e_J}\right]$ are, for all i ≠ J,

$$sign\left[\frac{\partial P}{\partial e_J}\right] = \begin{cases} sign\left[\frac{\partial P_J}{\partial e_J}\right] \; if \; [\sigma_{iJ} \; (1-\varepsilon_{ii} + \varepsilon_{ii}\frac{P_J}{P_i})] \quad 0 \\ \\ -sign\left[\frac{\partial P_J}{\partial e_J}\right] \; if \; [\sigma_{iJ} \; (1-\varepsilon_{ii} + \varepsilon_{ii}\frac{P_J}{P_i})] \quad < 0 \end{cases} . \quad (7)$$

From equation (7), the sufficient conditions, that the market price moves in the same (opposite) direction as country J's price, are addressed as in the following proposition.

*[**Proposition 2**] When the exchange rate between a producer (say country J) and the destination market temporarily changes;*

(i) the market price and country J's price move in the same direction

$$if \;\; \sigma_{iJ} \underset{(<)}{>} 0 \;\; and \;\; P_J \underset{(<)}{>} \frac{\varepsilon_{ii} - 1}{\varepsilon_{ii}} P_i \;\; and;$$

(ii) the market price and country J's price move in different directions

$$if \;\; \sigma_{iJ} \underset{(<)}{>} 0 \;\; and \;\; P_J \underset{(>)}{<} \frac{\varepsilon_{ii} - 1}{\varepsilon_{ii}} P_i \; .$$

This proposition indicates that even when the cross price elasticity is positive, if country J's price is significantly larger than other countries' prices initially, then it is the case that the (weighted average) market price actually decreases (increases) as country J's price increases (decreases). This is possible as the market share country J loses due to the increase in its price is so large that sufficiently large weights are now given to other countries that have substantially lower prices and take market share from country J.

All possible combinations available from [Proposition 2] are summarized in Table 1. While the threshold of σ_{iJ}, which determines whether country i's price moves in the same direction as country J's, is zero, the threshold of the second condition determines whether J's price is larger or smaller than the product of P_i and an inverse of the famous 'price mark-up' ($\frac{P_i}{MC_i} = \frac{\varepsilon_{ii}}{\varepsilon_{ii} - 1}$) for an imperfectly competing firm i. While we have four cases (sufficient conditions) which have the same or different signs for ERPT for country J and the US market,

<Table 1> The Direction of the Two Exchange Rate Pass-Through and Sufficient Conditions

	$\frac{P_J}{P_i} > \frac{\varepsilon_{ii} - 1}{\varepsilon_{ii}}$	$\frac{P_J}{P_i} < \frac{\varepsilon_{ii} - 1}{\varepsilon_{ii}}$
$\sigma_{iJ} \geq 0$	(I) Same	(II) Opposite
$\sigma_{iJ} < 0$	(III) Opposite	(IV) Same

an assumption is added, which makes our analysis simpler.

[Assumption 2] Prices of products from different countries are not significantly different. [5]

This assumption allows us to preclude Cases II and IV in Table 1, and to avoid tedious repetition. Case I occurs when J's price in the US decreases (increases) and the other countries respond by decreasing (increasing) their prices. Then, the market price will decrease, with Assumption 2. This is so, regardless whether country J responds *normally* (for example, to increase / decrease its price by appreciation / depreciation of its currency) or *perversely* (for example, to decrease / increase its price by appreciation / depreciation of its currency) to the fluctuation of the value of its currency. If other countries do not follow country J and respond by changing prices in the different direction, the market price P will move in the different direction from country J's price, with Assumption 2. This can be explained as the following; the decrease (increase) in P due to the decrease (increase) in P_i (and the increase in ω_i) is sufficiently large to offset the increase (decrease) in P by the increase (decrease) in P_J. This is summarized in Case III. Therefore, the inequality $P_J > \frac{\varepsilon_{ii} - 1}{\varepsilon_{ii}} P_i$ can be understood as the condition necessary to allocate a sufficiently large market share to the country with relatively lower prices, and, consequently, lead to a reduction of the weighted average market price.

Table 1 also presents the possibility of the *'J-curve'* for the market ERPT, which was theoretically suggested by Tivig (1996) for individual countries. More discussion on this issue will be done in Section III with empirical findings.

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[5] If demand is not very elastic, $[(\varepsilon_{ii} - 1)/\varepsilon_{ii}]$ is considerably smaller than one and it is unrealistic to have such a small P_J that $P_J < [(\varepsilon_{ii} - 1)/\varepsilon_{ii}]\ P_i$ as the products are quite homogenous. In contrast, if demand is very elastic, which indicates that the products are highly substitutable and prices must be very close to each other, then $[(\varepsilon_{ii} - 1)/\varepsilon_{ii}]$ will be close to one. In either case, it is more plausible to assume that $P_J > [(\varepsilon_{ii} - 1)/\varepsilon_{ii}]\ P_i$. In addition, when producers operate in the inelastic region $(\varepsilon_{ii} < 1)$, Cases II and IV are not available as $[(\varepsilon_{ii} - 1)/\varepsilon_{ii}]$ is negative.

III. Empirical Analysis of Exchange Rate Pass-Through

1. The US Steel Market and Data

We are interested in the steel market for several reasons. First of all, in the era of e-commerce and genomics, the importance of steel to the economy has never diminished. As reported by the World Steel Dynamics (1997, 2000), the world steel market has experienced repeated 'booms' and 'downturns'. Nevertheless, steel has continued to be regarded as one of the most important materials and inputs necessary for major economic activities.

The high volatility of steel prices, which is not removed even when real values are investigated, also makes research on steel prices significant and interesting. This volatility has increased over time, and even appeared to accelerate as the steel industry throughout the world entered an era of restructuring (World Steel Dynamics, 1997). While various factors have contributed to this volatility, ranging from technological innovations and computerized production planning and management, to excessive investment in the steel industry of many countries, the volatility of steel product prices may also, in part, be explained by the volatility of exchange rates (Tcha and Sjaastad, 1998).[6]

In particular, this study concentrates on five steel products, which are quite disaggregated and may be categorized into three groups based on the characteristics of the production process and the final use of the product. Group 1 consists of two steel products; hot-rolled strips and hot-rolled sheets, where hot-rolled strips are used as an input to produce hot-rolled sheets. Group 2 consists of another two steel products; hot bars and cold bars.[7] Group 3 comprises three steel products; hot-rolled strips, cold-rolled strips and cold-rolled sheets, which are related from the upstream product to the downstream product. Utilizing data for these six steel products is in many ways unique. Most of the previous empirical studies on ERPT observed oligopoly markets such as automobiles, where products are fairly differ-

[6] The volatility of steel prices may be also due to protection. While we recognize the relevance of this variable, it is not included in this study as accurate data for various forms of protection imposed by the US are not available, including the problem of converting quantitative restriction into tariff-equivalent measure of protection. Further investigation is needed to overcome this problem.

[7] It is technically controversial whether we can categorize specific steel products into 'hot bars' and 'cold bars'. However, the American Metal Bulletin regularly publishes the related data, such as prices and market shares for these categories of products. Accordingly, we follow the same approach and use the data published.

entiated (for example, Feenstra et al., 1996; Gross and Schmitt, 2000; Laussel et al., 1988) or other manufacturing (Fisher, 1996; Marston, 1990), or some commodity at the aggregated level (Varangis & Duncan, 1993). The steel products examined in this paper are much more disaggregated and almost homogenous in their quality and characteristics, regardless of where they are exported from. These features are consistent with the assumptions adopted in this study. Also, we hope that the market ERPT by the two exchange rates (Korean won–US dollar and Japanese yen–US dollar) will enable us to compare the individual ERPT of almost homogeneous products from Korea and Japan, and will reveal the difference between the two countries' pricing strategies, which were the focus of attention in the previous studies by Klitgaard (1999) and Kim (1997).

In this study, all steel prices are obtained from The Statistical Guide to the Metal Industries: Metal Statistics (various years), and the economic variables, such as exchange rates and petroleum prices, are collected from International Financial Statistics (various years). All prices and exchange rates are converted into real terms using consumer price indices. The data is quarterly from 1970:1 to 1996:4 comprising 108 observations. The data since 1997 was excluded form the analysis as a kind of structural changes in the market was expected due to the economic crisis and accompanied violent swing in Korea-US exchange rates, where considering the market with structural break is not the major concern of this study. For simplicity, we labelled the price of each steel product in the following manner: P_1 for hot-rolled strips, P_2 for hot-rolled sheets, P_3 for hot bars, P_4 for cold bars, P_5 for cold-rolled strips and P6 for cold-rolled sheets.

2. Estimation Methodology and Statistical Inference

If the model discussed in the previous section is generalized over multi-periods, each country will maximize its present discounted value of profit in its own currency, in the first period, based on Tivig (1996), Froot and Klemperer (1989) and Gross and Schmitt (2000). Country J's optimal price at t is, therefore, in a general form,

$$P_{Jt} = F(\mathbf{P}_J', \mathbf{P}_O', \mathbf{e}_J', \mathbf{e}_O', \mathbf{MC}_J),$$

where $\mathbf{P}_J$ is a vector of past prices of the steel product of country J, $\mathbf{P}_O$ is a matrix of past prices of the steel product from all the other countries, $\mathbf{e}_J$ is a vector of past and present exchange rates between country J's currency and US dollar, $\mathbf{e}_O$ is a vector of past and present exchange rates between each country's currency and US dollar, and $\mathbf{MC}_J$ is the past and present marginal cost of the product in country J.

Among the various resources, raw materials and labor required to produce steel products, our study adopts petroleum as an index that captures the effect on costs of price changes in raw materials, whose significance was confirmed by Tcha and Sjaastad (1998). We consider the vector autoregressive (VAR) model of the form

$$Y_t = B_1 Y_{t-1} + \ldots. + B_p Y_{t-p} + \delta t + u_t, \tag{8}$$

where Y_t is the K×1 vector of variables including relevant prices and exchange rates at time t, and B_i's are the K×K matrices of coefficients. The model also contains the intercept vector and linear time trend terms, however, these deterministic components do not appear in the VAR representation for simplicity of exposition, while they are included in estimation. Note that u_t is the K×1 vector of i.i.d. innovations with $E(u_t) = 0$ and $E(u_t u_t') = \Sigma_u = HH'$.

The above VAR system can be written in the vector error correction form as

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} + \Pi (Y_{t-1} + \gamma t) + u_t, \tag{9}$$

where $\Pi = B_1 + \ldots + B_p - I_K$, $\Gamma_i = -(B_{i+1} + \ldots + B_p)$ and $\delta = \Pi\gamma$. When Y_t is cointegrated with the cointegration rank r, $Rank(\Pi) = r < K$ and $\Pi = ab'$, where a and b are respectively K×r matrices.

The VAR model is fitted to each group of steel prices and explanatory variables. In each case, the vector Y_t contains e_J (say, hereafter the value of US dollar in terms of Japanese yen), e_K (hereafter the value of US dollar in terms of Korean won), and the petroleum price followed by steel prices from upstream to downstream products in the group. That is, five-dimensional VAR is fitted for Groups 1 (hot-rolled strips – hot-rolled sheets) and 2 (hot bars – cold bars), while a six-dimensional VAR is fitted for Group 3 (hot-rolled strips – cold-rolled strips – cold-rolled sheets). The ordering of the variables in Y_t is based on the Wold causality (Lütkepohl, 1991, p.52), which is important in the context of the impulse response analysis based on the VAR. It indicates that contemporaneous causality runs from e_J and e_K to steel prices, and not in the opposite direction, which is consistent with our intuition.

For each group, the VAR order in (8) is determined so as to ensure that the least-squares residuals of each equation in VAR mimic a white noise. We attempt to find the smallest order possible for parsimonious parameterization. To this end, a visual inspection of the residual autocorrelation function (ACF) is conducted, accompanied by the use of the Ljung-Box (1978) test. It is found that the VAR(4) is adequate for Groups 1 and 3, while the VAR(3) is adequate for Group 2. To test for the cointegration for each model we use the method developed by Johansen (1988). We also use the fully modified OLS estimator of Phillips and Hansen (1990) to estimate the long-run relationship among the variables.

Impulse response analysis is conducted to examine short-run dynamics among the variables. This form of analysis is also closely related to causality, as zero impulse responses between two variables means that no dynamic causality exists between them (Lütkepohl, 1991).[8]

3. Empirical Findings and Discussions

Non-stationarity of Data and Cointegration

The presence of non-stationarity for the time series variables suggests that a spurious regression problem may exist. Table 2 reports the augmented Dickey-Fuller (ADF) test statistics for all variables. Two ADF statistics are reported: one (τ_μ) from the regression with intercepts but with no trends, and the other (τ_τ) from the regression with intercepts and linear trends. The order of augmentation is determined using Akaike Information Criteria (AIC). For most cases, the null hy-

<Table 2> Unit-Root Tests: ADF Statistics

Variable	τ_μ	τ_τ
P1	-3.03 (4)	-1.37 (4)
P2	-2.52 (3)	-1.29 (3)
P3	-2.48 (0)	-2.39 (0)
P4	-2.98 (0)	-1.16 (0)
P5	-2.44 (0)	-1.84 (0)
P6	-2.53 (0)	-1.28 (0)
e_J	-1.78 (2)	-2.40 (2)
e_K	-1.85 (2)	-1.91 (2)
PE	-2.48 (0)	-2.86 (0)

Note: τ_μ is the ADF statistic based on the model with intercept but no trend.
τ_τ is the ADF statistic based on the model with intercept and trend.
The numbers in brackets next to the statistics are the order of augmentation chosen by the AIC.
The 5% critical values for τ_μ and τ_τ statistics are –2.89 and –3.36.
P1: hot-rolled strips, P2: hot-rolled sheets, P3: hot bars, P4: cold-rolled strips, P5: cold-rolled sheets, P6: cold bars, e_J: US-Japanese exchange rate, e_K: US-Korean exchange rate, PE: petroleum price

[8] More information on impulse response analysis using bootstrapping technique is discussed in Appendix.

pothesis of a unit root cannot be rejected at the 5% level of significance. One exception is the τ_μ statistic for P_1, where the null hypothesis is accepted. However, P_1 shows an upward trend, and, in this case, the use of the τ_τ statistic should be more appropriate. Note that there is no evidence of the second unit root for any time series. Hence, we concluded that all time series are integrated of order one. Table 3 reports the trace and maximum eigenvalue (λ_{max}) statistics of Johansen (1988) for each model. For Groups 1 and 2, the null hypothesis of no cointegrating vector in favor of at least one cointegrating vector is rejected, but the possibility of three cointegrating vectors cannot be rejected at the 5% level of significance. For Group 3, the λ_{max} statistic indicates acceptance of two cointegrating vectors, but the statistic is fairly close to the 10% critical value. Hence, it seems reasonable to conclude that Group 3 is cointegrated with three cointegrating vectors.

Vector Error Correction Models and Impulse Response

Table 4 reports the estimated error correction models for the price variables. For each group, where error correction terms were obtained

<Table 3> Test Statistics for Cointegration

Group 1

Null	Trace	Trace (0.95)	λ_{max}	λ_{max}(0.95)
r = 0	129.03	87.17	61.93	37.86
r ≤ 1	67.10	63.00	29.21*	31.79
r ≤ 2	37.89	42.34	18.45	25.42
r ≤ 3	19.44	25.77	14.08	19.22
r ≤ 4	5.36	12.39	5.36	12.39

Note: *: Significant at the 10% level of significance
Trace (0.95) and λ_{max}(0.95) indicate 5% critical values for each statistic.

Group 2

Null	Trace	Trace (0.95)	λ_{max}	λ_{max}(0.95)
r = 0	117.26	87.17	48.31	37.86
r ≤ 1	68.95	63.00	29.15*	31.79
r ≤ 2	39.80	42.34	18.96	25.42
r ≤ 3	20.84	25.77	11.94	19.22
r ≤ 4	8.90	12.39	8.90	12.39

Note: *: Significant at the 10% level of significance
Trace (0.95) and λ_{max} (0.95) indicate 5% critical values for each statistic.

<Table 3> Continued
Group 3

Null	Trace	Trace (0.95)	λ_{max}	λ_{max}(0.95)
r = 0	183.15	115.85	64.21	43.61
r ≤ 1	118.94	87.17	55.88	37.86
r ≤ 2	63.06	63.00	28.99	31.79
r ≤ 3	34.07	42.34	20.01	25.42
r ≤ 4	14.06	25.77	8.56	19.22
r ≤ 5	5.50	12.39	5.50	12.39

Note: Trace (0.95) and λ_{max}(0.95) indicate 5% critical values for each statistic.

using Johansens's (1988) just identifying restrictions, the vector error correction (VEC) models associated with the cointegrating regressions are estimated. All error correction models show a reasonably good fit with no sign of model mis-specification, including serial correlation in error terms. This is evident from visual inspection of the residual ACF and Ljung-Box test statistics as well as the Durbin-Watson statistics reported. For all models, it can be seen that the current price changes are to some extent affected by the short-run changes in exchange rates and factors affecting marginal cost such as petroleum prices (PE). As the results associated with these VEC models are sometimes unclear and hard to interpret, we next use impulse response analysis to examine the short-run dynamics of the variables involved. These results are summarized in Figures 1 to 3.

Figure 1 plots impulse response functions of P_1 and P_2 in Group 1 against time horizon 0 to 24 (0 to 24 quarters), when one standard deviation shock is given to e_J, e_K, PE, P_1 and P_2. Bootstrap-after-bootstrap confidence intervals with a probability content of 90% and 95% are given for statistical inference. If a 95% (90%) confidence interval contains zero, the null hypothesis of zero impulse response value cannot be rejected at the 5% (10%) level of significance. It can be seen that e_J shows no dynamic impact on P_1 and P_2, as all confidence intervals contain zero. This result implies on the one hand, that there is no direct and clear evidence that changes in e_J affect the dollar price of the Japanese steel products, such as hot-rolled strips and sheets (and therefore the weighted average market price does not change). Alternatively, with the given data described in the previous section, it might be concluded that while the price of Japanese steel products are affected by the exchange rate (i.e., $[\partial P_J/\partial e_J] \neq 0$), the weighted average market price of the products are not significantly affected by the

<Table 4> Exchange Rate Pass-Through: Error-Correction Models

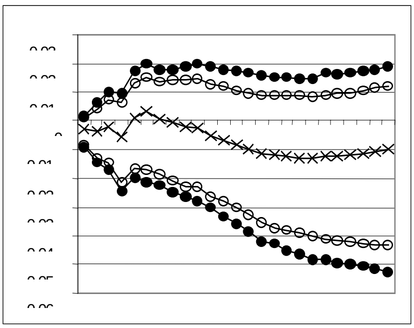
	Group 1		Group 2		Group 3		
	ΔP1	ΔP2	ΔP3	ΔP5	ΔP1	ΔP4	ΔP6
ecm1	0.10***	-0.01	0.26***	0.13**	-0.06**	-0.07*	-0.03
ecm2	-0.05*	0.11*	-0.08	0.22**	-0.11***	0.04	-0.13***
ecm3					0.07***	0.07*	-0.00
ΔP1(-1)	-0.08	0.01			-0.07	-0.17	0.51**
ΔP1(-2)	0.21**	0.65***			0.23*	0.08	0.38*
ΔP1(-3)	0.18*	0.31			0.22**	0.17	0.20
ΔP2(-1)	-0.10	-0.02					
ΔP2(-2)	-0.06	-0.37***					
ΔP2(-3)	-0.05	-0.12					
ΔP3(-1)			0.01	-0.01			
ΔP3(-2)			0.21*	-0.06			
ΔP3(-3)							
ΔP4(-1)					-0.18*	0.05	-0.18
ΔP4(-2)					-0.13	-0.06	-0.13
ΔP4(-3)					-0.05	0.12	-0.12
ΔP5(-1)			0.04	0.09			
ΔP5(-2)			-0.21	-0.01			
ΔP5(-3)							
ΔP6(-1)					0.15**	-0.01	0.03
ΔP6(-2)					0.10	-0.05	-0.09
ΔP6(-3)					0.01	-0.21**	-0.05
Δe_J (-1)	-0.13**	-0.20	-0.33**	-0.23**	-0.03	-0.07	0.03
Δe_J (-2)	-0.09	0.05	0.13	0.04	-0.01	-0.04	0.09
Δe_J (-3)	-0.15**	0.20			-0.10	-0.08	0.04
Δe_K (-1)	0.37***	0.19	-0.49**	0.22	0.27**	0.10	-0.25
Δe_K (-2)	-0.57***	-0.84***	-1.03***	-0.42**	-0.48***	-0.16	-0.48***
Δe_K (-3)	-0.18	-0.16			-0.18	-0.23	0.03
ΔPE(-1)	0.02	0.03	0.00	-0.03	-0.02	0.04	-0.03
ΔPE(-2)	0.01	0.03	0.00	-0.08***	-0.01	-0.02	0.00
ΔPE(-3)	0.02	0.00			-0.00	0.01	-0.20
R^2	0.50	0.36	0.30	0.30	0.62	0.31	0.22
DW	2.14	1.90	2.07	2.12	2.02	2.05	1.85

Note: *** : Significant at the 1% level of significance
** : Significant at the 5% level of significance
* : Significant at the 10% level of significance
DW : the Durbin-Watson statistic
P1 : hot-rolled strips, P2: hot-rolled sheets, P3: hot bars, P4: cold-rolled strips, P5: cold-rolled sheets, P6: cold bars, e_J: Japan-US exchange rate, e_K: Korea-US exchange rate, PE: petroleum price

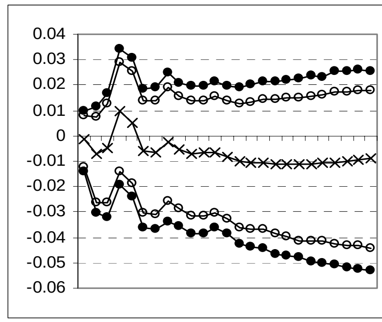
[Figure 1] Orthogonalized Impulse Response Estimates and Confidence Intervals: Group 1

Shock to e_J

P1

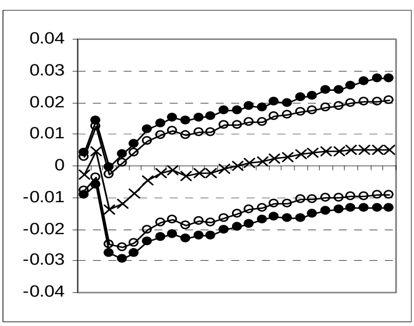


P2

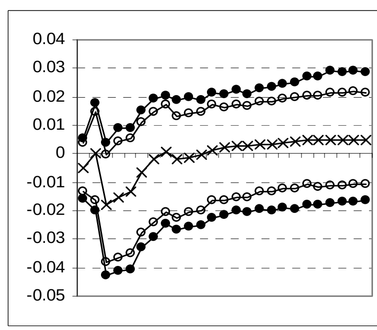


Shock to e_K

P1

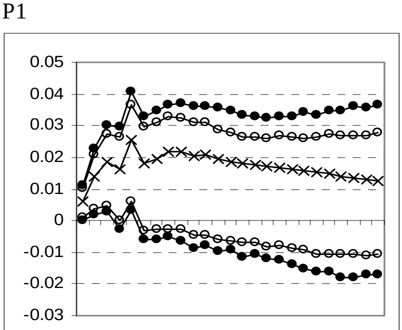


P2

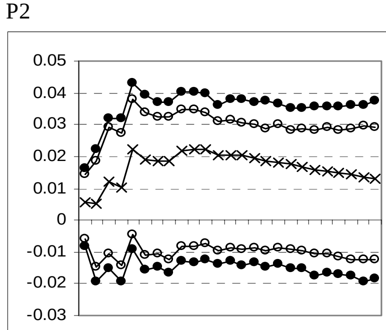


Shock to PE

P1



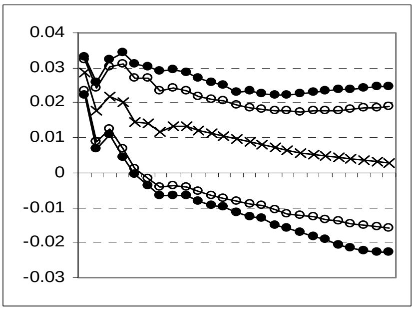
P2



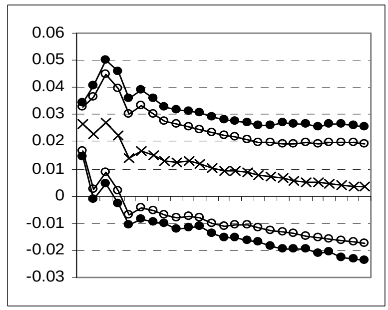
[Figure 1] Continued

Shock to P1

P1

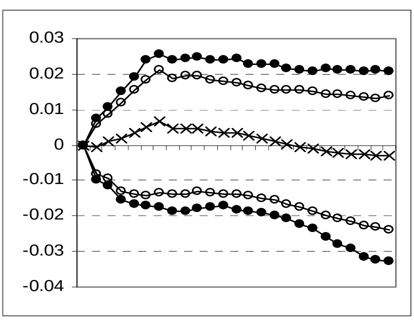


P2

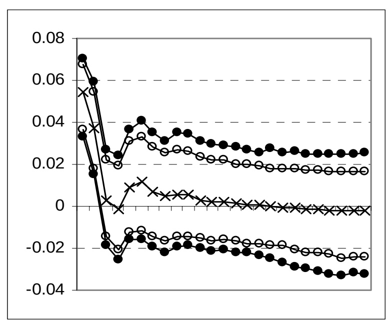


Shock to P2

P1



P2



Note ● : 95% Bootstrap-after-bootstrap Confidence Intervals
○ : 90% Bootstrap-after-bootstrap Confidence Intervals
× : Impulse-response estimates
The X-axes of all graphs indicate time horizon from 0 to 24

change in Japanese prices (i.e. $\left[1-\ \phi_{JJ}\ +\sum_{i\,\neq\,J}(1\ -\ \phi_{ii})\ \sigma_{iJ}\ \frac{\omega_i P_i}{\omega_J P_J}\right]$ from equation (5) is insignificantly different from zero).

In contrast, changes in e_K 'pass through' and change the dollar price of the commodities in the US market, with lags. There is some evidence that e_K affects P_1 and P_2 negatively (*normal market ERPT*) with a time lag, as 95% and 90% confidence intervals do not contain zero at the second quarter after the exchange rate shock, respectively for P_1 and P_2. The price of petroleum (*PE*) affects the price of hot-rolled strips, P_1, positively for five quarters once the shock is given, but exerts no impact on the price of the commodity at the end of the production process, hot-rolled sheets (P_2). This result is plausible considering that the impact of the petroleum price on the final product weakens as changes in the petroleum prices partially affect the price of hot-rolled strips, where, in turn, the price of hot-rolled strips only partially explains the price of hot-rolled sheets. It is also evident that P_1, the price of hot-rolled strips, affects P_2, the price of hot-rolled sheets, for two quarters, but that P_2 has no impact on P_1. Both P_1 and P_2, depend on their own pasts for several quarters.

Figure 2 presents the impulse responses in Group 2. A shock in e_J has no impact on P_3, the price of hot bars, and P_4, the price of cold bars; but a shock in e_K affects P_3 negatively (*normal market ERPT*) after two and three quarters. P_4 is also negatively (*normal market ERPT*) affected by the shock to the Korean-US exchange rate for four quarters after the shock is given, although the impact is very marginal. The price of petroleum shows no impact on P_3. It exerts a positive impact on P_4 from quarters 4 to 7, however, the effect is marginal. It is again evident that changes in P_3 dynamically cause changes in P_4, but there is no evidence of causality from P_4 to P_3. Both Groups 1 and 2 report that the impulse response of the prices of downstream products (cold-rolled strips and cold bars) to the shock in the prices of upstream products (hot-rolled strips and hot bars) is significant; but the reverse is not the case.

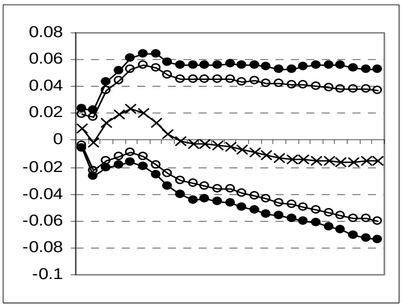
Figure 3 presents Group 3, and looks at the production stages of hot-rolled strips, cold-rolled strips, and cold-rolled sheets. Again, a shock in e_J shows no impact on P_1 and P_6. There is some evidence that e_J affects P_5 negatively; however, this has little practical implication as it happens after 12 quarters and the level of significance is very marginal. In contrast, e_K affects steel prices negatively (*normal market ERPT*) in the relatively short-term, with a lag of two or three quarters. The price of petroleum (*PE*) shows a positive impact on P_1 for four quarters. Some evidence indicates that it affects P_5 positively in the short-run, although less severely than it affects P_1, but not P_6. Among prices, causality runs from P_1 to P_5 and then to P_6; however, little evidence of causality in the opposite direction was found.

From the evidence related to the short-run dynamics adopting bootstrap-after-bootstrap methods for confidence intervals, the following general features emerge. First, a shock in the Japanese-US exchange rate in general shows no impact on steel prices in the US market (except in the case of P_5 in Group 3, where the response to impulse is found after three years), while a shock in the Korean-US exchange rate affects them negatively (normal market PT) with a lag of two or three quarters. However, these results do not directly support Klitgaard (1999) and Kim's (1997) assertion that Japanese firms actively absorbed the change in exchange rates by modifying

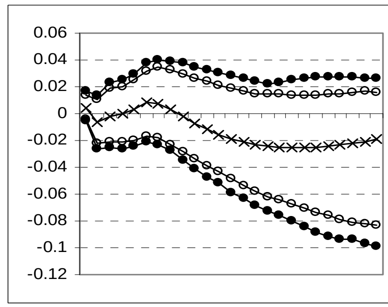
[Figure 2] Orthogonalized Impulse Response Estimates and Confidence Intervals: Group 2

Shock to e_J

P3

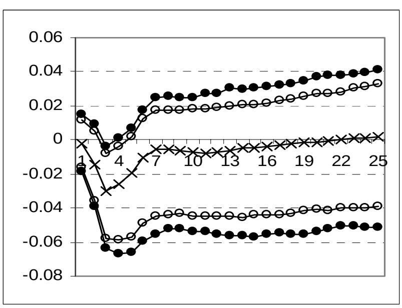


P5

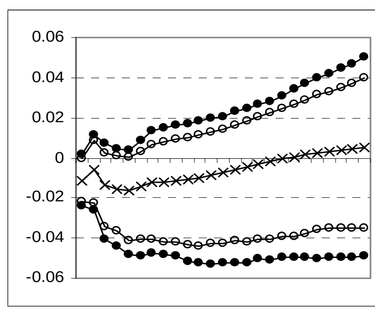


Shock to e_K

P3

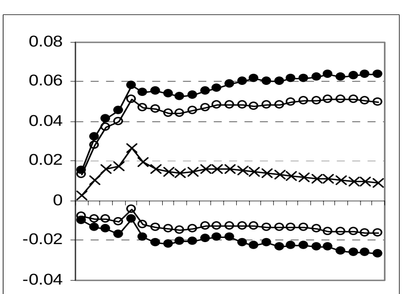


P5

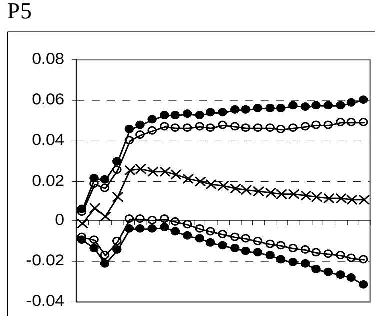


Shock to PE

P3



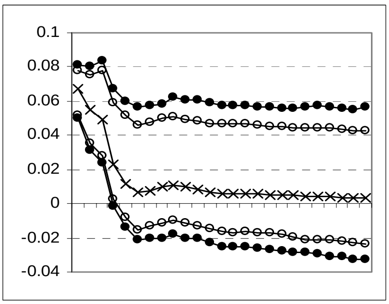
P5



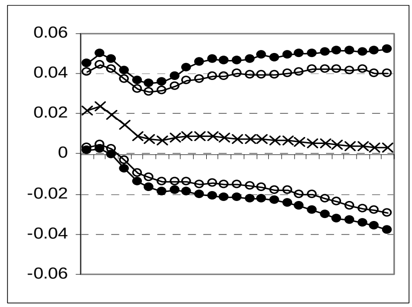
[Figure 2] Continued

Shock to P3

P3

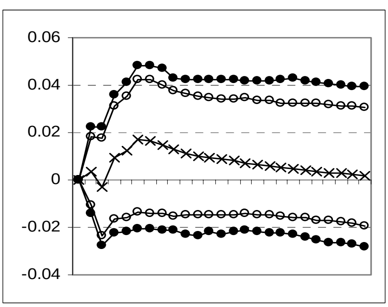


P5

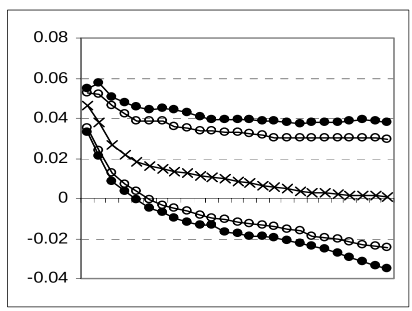


Shock to P5

P3



P5

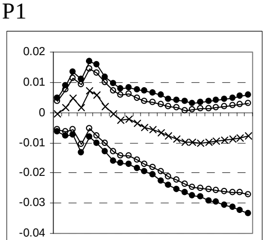


Note ● : 95% Bootstrap-after-bootstrap Confidence Intervals
○ : 90% Bootstrap-after-bootstrap Confidence Intervals
× : Impulse-response estimates
The X-axes of all graphs indicate time horizon from 0 to 24

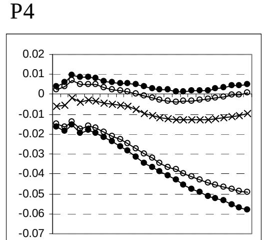
[Figure 3] Orthogonalized Impulse Response Estimates and Confidence Intervals: Group 3

Shock to e_J

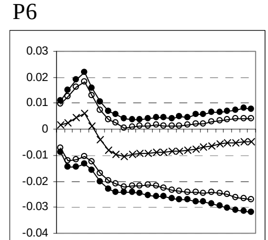
P1



P4



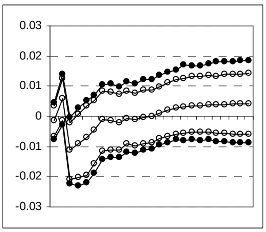
P6



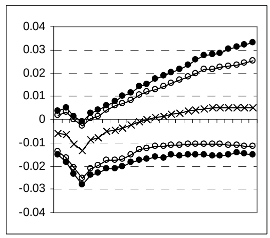
[Figure 3] Continued

Shock to e_K

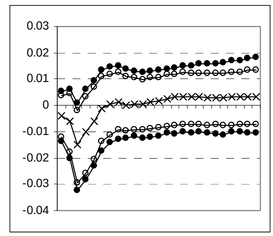
P1



P4

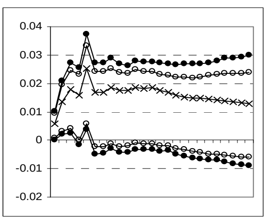


P6

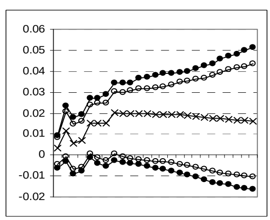


Shock to PE

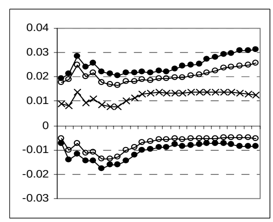
P1



P4

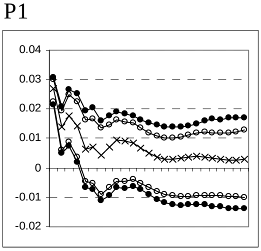


P6

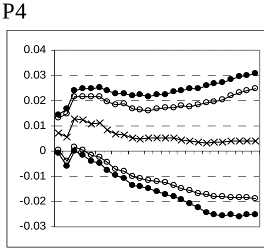


Shock to P1

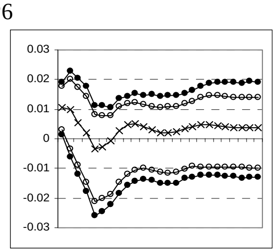
P1



P4

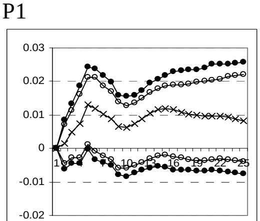


P6

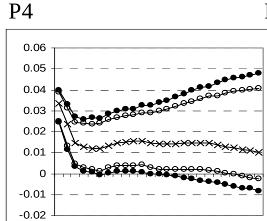


Shock to P4

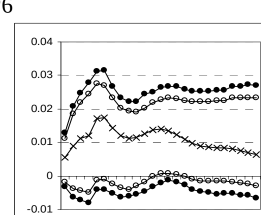
P1



P4



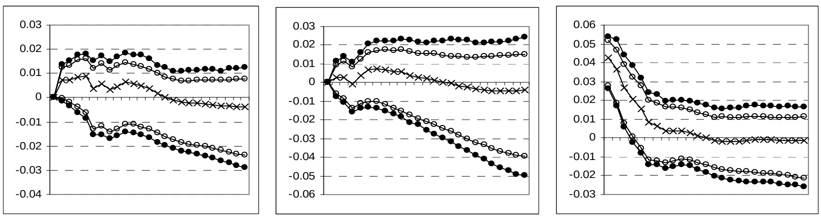
P6



[Figure 3] Continued

Shock to P6

P1 P4 P6



N*ote* ● : 95% Bootstrap-after-bootstrap Confidence Intervals
○ : 90% Bootstrap-after-bootstrap Confidence Intervals
× : Impulse-response estimates

The X-axes of all graphs indicate time horizon from 0 to 24

their profit margins or mark-ups while Korean firms did not. It is still possible to interpret our results that while the *individual ERPT* happens to Japanese producers, there is no *market ERPT*.[9]

In contrast to the Japanese case, Korean steel producers let the pass-through happen, at least partially. Any change in petroleum price, which represents marginal cost conditions, affects steel prices positively in the short-run, but there is a strong tendency that it is more likely to affect those products, which are in the upstream stages of production. The more processed the product is, the less affected it is by the price of petroleum. There is little evidence that the petroleum price affects the prices of the final products. Utilizing commodities in the same production stream, the findings of this study break new ground and go further than previous studies, which concentrated on the final goods in different or differentiated categories (for example, Feenstra et al., 1996; Fisher, 1996; Gagnon & Knetter, 1995; Laussel et al., 1988). For the prices of the steel products in the same production stream, causality runs from the prices of upstream products to the prices of downstream or final products, but not in the opposite direction.

Market ERPT, Individual ERPT and the J-curve

Market ERPT against the two exchange rate shocks shows *'J-curve'* cases in Figures 1 to 3.[10] The US market prices initially responded

[9] Another possibility to explain this phenomenon of no-response to Japanese price is the existence of the switching costs as illustrated by Klemperer (1995), Tivig (1996) and Gross and Schmitt (2000).

[10] The possibility of experiencing the 'J-curve' for individual ERPT under certain conditions was discussed in Tivig (1996). In that paper, it was shown that the J-curve in trade balance

positively to the fluctuation in e_J, although only insignificantly, and in turn dropped negatively, which is contrary to its response to the fluctuation in e_K. While Tivig (1996) entertains the theoretical possibility of a J-curve for an individual producer's ERPT[11], this J-curve has not been thoroughly discussed in empirical works, although many studies reported that the prices of specific goods from some countries moved perversely when its exchange rate changed. Moreover, there has been neither any theoretical nor empirical work done for the J-curve in association with the market ERPT.

If we have two firms competing in the market over two periods, each firm's strategic pricing behavior against exchange rate shocks results in 16 different cases.[12] With assumptions that capital is imperfectly mobile over time, and that producers choose a closed-loop solution where they can observe and respond to their opponents' actions at the end of each period, Tivig (1996) shows that second period price movements of each producer occur in opposite directions.[13] It is also impossible for country *J* to respond perversely for both periods if it maximizes intertemporal profit. These two findings remove 10 cases out of 16, and Table 5 summarizes all the remaining possible cases for individual ERPT. The last row for each case in Table 5 indicates the direction of the US market price movement for each case of the individual ERPT. Figures 1 to 3 indicate that the market ERPT against e_J shows similar movements to Case 5, while that against e_K demonstrates similar movements to Cases 2 and 4.

Case 5 for Japan explains that when the Japanese currency depreciates ERPT for Japan is likely to be normal (Japanese price decreases in the US market) in the first period, and then change to be perverse (Japanese price increases) during the next period, although only insignificantly.[14] All the other countries respond to Japan's price change by increasing their prices initially and then decreasing them. Cases 2 and 4 respectively, explain that Korean producers respond to the exchange rate shock either perversely (Case 2) or normally (Case 4) in

can be observed due to the perverse movement of individual price in the first period, in a case of a single good.

[11] According to his two-period model, all the other countries should increase (decrease) their prices in the second period responding to country *J*'s price increase (decrease) in the first period. By strategically behaving like this, firms can increase their market share over time and then compensate for the losses incurred by 'perverse' movement in the first period.

[12] Two firms' pricing behavior (price increases or decreases) over two periods with two different results in market average price (market price increases or decreases) provides 2*2*2*2 = 16 cases.

[13] In order to derive these findings, Tivig (1996) adopted more assumptions such as that demand functions are linear and that second-period demand functions are taken to be linear in the first period market shares.

[14] The argument in this part uses sufficient conditions as discussed in the previous sections, and hence, we cannot exclude the possibility that the market ERPT occurs due to some other reasons from those discussed here.

<Table 5> Directions of Market ERPT and Individual ERPT Over Time

Case	1st period	2nd period
1 P_J	↑	↓
P_i	↑ (+)	↑ (−)
P	↑	↑
2 P_J	↑	↓
P_i	↓ (−)	↑ (−)
P	↓	↑
3 P_J	↓	↑
P_i	↓ (+)	↓ (−)
P	↓	↓
4 P_J	↓	↓
P_i	↓ (+)	↑ (−)
P	↓	↑
5 P_J	↓	↑
P_i	↑ (−)	↓ (−)
P	↑	↓
6 P_J	↓	↓
P_i	↑ (−)	↑ (−)
P	↑	↑

Note: ↑ and ↓ mean 'increase' and 'decrease'. The signs in parentheses are those for σ_{iJ}. P_J and P_i are country J's and i's prices respectively, and P is market price.

the first period, however, only normally in the second period. It is possible for Korea to increase its prices temporarily in the first period in case of a depreciation of its currency, if its first-period market demand is inelastic, which is Case 4. However, it should reduce its prices in the second period to regain the market share that it lost due to the perverse movement in the first period. In contrast, all of the other producers decrease their prices in the first period regardless of Korea's strategy, take more market share, and can increase prices in the next period.

In addition, Case 4 for Korea indicates that the market J-curve can be observed even when the country's individual ERPT is normal over the periods. In this case, the main force to increase price in the second period is the increase in the price of the other countries after they retain large market shares in the first period by moving together with Korea (that is to decrease prices), and then increase their prices in the second period if their demand is inelastic. This J-curve for the market price is not found when the Japan–US exchange rate fluctuates.

Long-run Elasticities: Phillips-Hansen Fully Modified OLS

For each group, fully modified OLS estimation of Phillips and Hansen (1990) was conducted to examine the long-run elasticities among the variables, taking each price variable as a dependent variable. Table 6 summarizes the results obtained, which reveal that in the long run most of the elasticity estimates among the price variables are, when statistically significant, close to one. It should be pointed out, that long-run elasticities are found in most cases to be significant in both directions of the production line. In the long-run, changes in the price for any steel product are almost equi-proportionally transferred to those of other steel products in the same production line, regardless of whether they are in the upstream or downstream production line.

The elasticities of the US market price with respect to the Japanese-US exchange rate, in the long-run were statistically significant in four cases out of seven. Excluding the case of P_2 in Group 1 where the level of significance is marginal, the market price responds negatively (which means that the US market price decreases) in the long-run as the Japanese yen depreciates against the US dollar. In the previous section, we observed that ERPT for the US market (and Japanese) prices against e_J were initially negative (normal), and then positive(perverse). This implies that, while impulse response was, in general, insignificant, the perverse movement of the market and Japanese prices in the later periods is not sufficient to offset their initial normal movement. This is completely contrary to the findings of other studies, which argued that ERPT infrequently occurs normally to Japanese products.

<Table 6> Estimates of Long-Run Elasticities: Fully Modified OLS Estimation

	Group 1		Group 2		Group 3		
	P1	P2	P3	P5	P1	P4	P6
P1		0.97***				1.04***	0.84***
P2	0.93***						
P3				0.78***			
P4					0.59***		0.13
P5			0.93***				
P6					0.25***	0.17	
e_J	- 0.21***	0.10*	- 0.24***	- 0.04	- 0.06*	0.05	- 0.05
e_K	0.16*	- 0.06	0.45**	- 0.16	0.26***	- 0.30***	- 0.12
PE	0.01	0.01	- 0.06	0.14***	0.02*	0.01	- 0.02

Note: *** : Significant at the 1% level of significance
** : Significant at the 5% level of significance
* : Significant at the 10% level of significance
P1 : hot-rolled strips, P2: hot-rolled sheets, P3: hot bars, P4: cold-rolled strips, P5 : cold-rolled sheets, P6: cold bars, e_J: Japan-US exchange rate, e_K: Korea-US exchange rate, PE : petroleum price

These striking results are also found when the elasticities with respect to the Korean-US exchange rate were considered. The elasticities are significant in four cases out of seven, including one marginal case. While the depreciation of the Korean won decreases P_5 in Group 3 as traditional pass-through models expect, it increases market prices expressed in dollars for P_3 in Group 2 and for P_1 in Groups 1 and 3. Comparing the three products with significant elasticities, with respect to both e_J and e_K, provides an even more dramatic contrast: the price elasticities with respect to e_J are all negative, while those with respect to e_K are all positive. In other words, in the long-run, the depreciation of the Japanese yen against the US dollar decreases the weighted average prices of the products in the US market, while the depreciation of the Korean won against the US dollar increases the prices of the weighted average prices of the products in the US market. This implies that either perverse ERPT to Korean products in the initial period is so large that it dominates its normal ERPT in the later period, and/or an increase in the other countries' prices in the later period dominate. Alternatively, while the ERPT for Korea is normal over the periods, all the other countries' perverse reaction in the second period is so large that the market price is eventually pushed up.

Notice also, that e_J and e_K tend to affect the price of upstream products. The price for downstream products (which are close to the end user) is not so greatly affected by exchange rate fluctuations as that for upstream products. The price elasticities with respect to PE are statistically significant for P_4 and P_1, and the coefficients are positive, indicating that the prices of selected steel products respond positively to changes in petroleum prices.

IV. Summary

The effect of exchange rate fluctuations on the price of commodities has been explored by a host of studies due to its importance in trade balance and macroeconomic context (among many, see for example, Fitoussi & Cacheux, 1988; Froot & Klemperer, 1989; Krugman, 1989) and its importance in understanding competition in imperfect international markets (for example, see Feenstra et al., 1996; Tivig, 1996; and Gross and Schmitt, 2000). Unlike previous studies, this study concentrates on *the market ERPT* for similar commodities based on the theoretical pursuit of Tivig (1996) for *the individual ERPT*. While Sjaastad (1985), and Sjaastad and Scacciavilani (1996) explored the role of exchange rates on commodity prices, our study is unique in that it explores the mechanism of the market ERPT and the relationship between the individual ERPT and the market ERPT. The mechanism that market ERPT takes place depends on three stages: (i) pass-through by the exporter whose currency has changed in value, (ii) price responses by competitors of the exporter and (iii) changes in market share. From the market ERPT, the individual ERPT was found for two countries competing in the US market (Korea and Japan) as their different pricing strategies are of interest to economists and policy analysts.

Error correction models and impulse response analyses based on VAR contrast the different pricing strategies of Korea and Japan, or contrast the different magnitude of impact to the market. Within one or two years, changes in the Japanese-US exchange rates show no impact on the US market price, while changes in the Korean-US exchange rates are found to be significant. The response of the market price towards the fluctuation of both exchange rates empirically reveals the J-curve movement, which has not been previously explored at the market level.

The prices of upstream products were also affected significantly. However, these results are contrary to the conventional belief in previous studies once longer run effects are taken into account. For all products where fluctuations of both yen-dollar and won-dollar exchange rates are significantly influential, the depreciation of the Japanese yen against the US dollar, in fact, decreases the average dollar price of the product, while the depreciation of the Korean won against the US dollar increases it. This finding implies that the previous argument may not be valid in the long-run, as Japanese producers absorb the impact of exchange rate fluctuations by adjusting their profit or price mark-ups. In contrast, the argument that Korean steel producers aggressively change their prices in the US market as exchange rates change is only partially plausible; their long-run responses are frequently perverse.

In summary, our results show that (i) the market price response in the short-run can be different from that in the long-run, (ii) ERPT for the market price is not necessarily the same as ERPT for the country experiencing currency depreciation or appreciation, and (iii) countries may have different pricing strategies against exchange rate shocks even when products are similar.

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Appendix

Given m realizations $(Y_1, \ldots, Y_m)$ of (8), the unknown coefficients are estimated using the least-squares (LS) method. The LS estimators for $B = (B_1, \ldots, B_p)$ and Σ_u are denoted as $\hat{B} = (\hat{B}_1, \cdots, \hat{B}_p)$ and $\hat{\Sigma}_u$. The orthogonalized impulse responses are defined as $\Theta_r = \Phi_r H$, where $\Sigma_u = HH'$ and Φ_r's are the coefficients of the MA(∞) representation of (8). A typical element of Θ_r is denoted as $\theta_{kl,r}$, and is interpreted as the response of the variable k to a one-time impulse in variable l, r period ago. Using $\hat{B}$ and $\hat{\Sigma}_u$, the estimator for impulse response $\hat{\theta}_{kl,r}$ for $\theta_{kl,r}$ can be calculated.

The bootstrap-after-bootstrap confidence interval for $\theta_{kl,r}$ can be constructed as the following two stages. In Stage 1, we generate a pseudo data set following the recursion

$$Y^*_t = \hat{B}_1 Y^*_{t-1} + \ldots + \hat{B}_p Y^*_{t-p} + u^*_t , \tag{A1}$$

using the first p values of the original data as starting values. Note that u_t^* is a random draw from the residuals of the LS estimation of the model. Using $\{Y_t^*\}_{t=1}^{m}$, the coefficient matrices are re-estimated and denoted as $\hat{B}^* = (\hat{B}_1^*, \ldots, \hat{B}_p^*)$. Repeat this process to generate 5000 sets of $\{Y_t^*\}_{t=1}^{m}$ and obtain the corresponding 5000 bootstrap replicates of $\hat{B}^*$. The bias of $\hat{B}$ can be estimated as $\hat{\Psi} = \overline{B}^* - \hat{B}$, where $\overline{B}^*$ is the sample mean of 5000 bootstrap replicates of $\hat{B}^*$. The bias-corrected estimate for B can be obtained as $\tilde{B} = \hat{B} - \hat{\Psi}$, implementing the sta tionarity correction detailed in Kilian (1998)[1].

In Stage 2, generate a pseudo data set following the recursion

$$Y^*_t = \tilde{B}_1 Y^*_{t-1} + \ldots + \tilde{B}_p Y^*_{t-p} + u^*_t ,$$

in the same way as in (A1). Note, that we now use the bias-corrected parameter estimators to generate the pseudo data set $\{Y_t^*\}_{t=1}^{m}$. Re-estimate the coefficient matrices using $\{Y_t^*\}_{t=1}^{m}$ and the parameter estimator is denoted as $\hat{B}^*$. The bias-corrected estimator is obtained as $\tilde{B}^* = \hat{B}^* - \hat{\Psi}$, again implementing the stationarity cor-

[1] The stationary correction is given to prevent the parameter estimators from becoming non-stationary as a result of bias-correction.

rection. Repeat this process to generate 5000 bootstrap replicates of $\tilde{B}^*$, from which 5000 bootstrap replicates $\hat{\theta}^*_{kl,r}$ of impulse responses are obtained.

The 100(1-2α)% bootstrap-after-bootstrap confidence intervals for $\theta_{kl,r}$ can be constructed as the interval $[\hat{\theta}^*_{kl,r}(\alpha), \hat{\theta}^*_{kl,r}(1-\alpha)]$, where $\hat{\theta}^*_{kl,r}(g)$ is the gth percentile from the distribution of 5000 bootstrap replicates of $\hat{\theta}^*_{kl,r}$ based on the percentile method of Efron and Tibshirani (1993, p.160). This confidence interval can be used to test for statistical significance of impulse response estimates. If a 95% (90%) confidence interval contains zero, the null hypothesis of zero impulse response value cannot be rejected at the 5% (10%) level of significance. According to Killian (1998), this bootstrap-after-bootstrap confidence interval performs substantially better than other conventional alternatives for statistical inference of impulse response estimates, especially for the cointegrated VAR model when the sample size is small. Therefore, we firmly believe that the application of this method will certainly provide more robust and reliable results.