Essays on Applied International Economics and Finance

by

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Abstract

Studies in international economics and finance have become increasingly important in understanding the ways economies are integrated. This dissertation comprises three essays on this topic. The first essay investigates the relationship between financial development and trade based on panel data of bilateral trade between the world's three largest economies (United States, Japan, and Germany) and 47 partner countries. The second essay estimates common factors from a monthly panel of 51 commodity prices and further analyzes the most important common factor, which appears to be correlated with the U.S. dollar nominal exchange rates. Lastly, the third essay examines data on Japanese and Korean automobile exports to the United States to determine consistency with the Alchian-Allen theorem.

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CHAPTER 1

Financial Development and Trade:

Evidence from the World's Three Largest Economies

1.1 Introduction

Numerous studies examine the relationship between international trade and economic growth, as well as between financial development and economic growth.[1] The importance of trade and financial development in the growth literature provides motivation to study the relationship between the two. Many of the existing studies on this issue examine individually specific measures of financial development, e.g., private credit and foreign investment.[2] The present study aims to contribute to the literature by investigating several measures of financial development with a gravity model of bilateral trade between the world's three largest economies (United States, Japan, and Germany) and their 47 major trading partners from 2003 to 2007.

This paper examines access to external funds and international financial indicators as the measures of financial development. The main hypothesis tested is that financial development in a country relates to the degree of bilateral trade with its trading partners. A gravity model is constructed and estimated with a fixed effects method. The model includes three variables commonly used in a gravity equation: distance between pair countries, land common border, and stage of development. Ease of access to loans and ease of access to the local equity market represent access to external funds. Three international financial indicators are included: country credit ratings, international capital market controls, and real exchange rates. A country's credit rating and capital controls capture its access to foreign capital, while the real exchange rate plays a role in determining profitability of tradables.

[1] Frankel and Romer (1999), among others, find a positive relationship between international trade and economic growth. Several studies suggesting the importance of financial development for economic growth are Levine and Zervos (1998), Levine, Loayza, and Beck (2000), and Demirgüç-Kunt and Levine (2001).

[2] The World Economic Forum (2009, p. 3) defines financial development as "the factors, policies, and institutions that lead to effective financial intermediation and markets, as well as deep and broad access to capital and financial services."

Throughout this paper, the term ‘main countries’ will refer to the three largest economies, and ‘partner countries’ will refer to the 47 trading partners. The study finds differences in the degree of bilateral trade when grouping partner countries into developed and less developed countries. Access to loans for businesses shows a strong positive relationship with bilateral trade. Access to the local equity market is negatively related to trade with developed countries, but positively related to trade with less developed countries. The study also finds that country credit ratings, international capital market controls, and real exchange rates are significant determinants of trade. The study suggests the roles policy can play in promoting both trade and development.

The remainder of the paper proceeds as follows. Section 2 describes the data. Section 3 provides literature review. Section 4 and Section 5 discuss the empirical methods and results. Section 6 concludes the paper with a summary of findings.

1.2 Data

The dataset is a balanced panel of bilateral trade between the three main countries (United States, Japan, and Germany) and 47 partner countries over the years 2003 to 2007 for a total of 690 observations. Sample selection is based on the average levels of GDP of IMF reporting countries over the five year period and availability of other data. The top three countries based on the average GDP are selected as the main countries and the rest as partner countries, shown in Table 1.1. The total trade between the main countries and partner countries during the period represents about 70% of total trade of the three main countries.

[Table 1.1 here]

The major data source in this study is the International Monetary Fund (IMF) databases. The bilateral trade data are obtained from the IMF's Direction of Trade (DOT) database, the GDP data are from the IMF's World Economic Outlook (WEO) database, and the real effective exchange rate data are from the IMF's International Financial Statistics (IFS) database. The IMF trade data follows United Nation's guidelines, which sufficiently covers all merchandise entering or leaving a country, except goods in transit. The IFS database provides an indicator of real effective exchange rates based on relative consumer prices, allowing for comparison with a broad range of partner countries. The weighting method is based on disaggregated trade flows for manufactured goods and primary products over the period 1999 to 2001 (IMF, 2009).

Another data source is the World Economic Forum's (WEF) Global Competitiveness Reports (2003-2007). The WEF reports the indices of access to loans (Ease of Access to Loan) and access to the local equity markets (Ease of Access to Equity). These indices are scaled 1 to 7 with higher scale indicating easier access to obtain loans or to raise capital through the local equity market. They are based on surveys by the WEF on the opinions of business executives worldwide. The same report contains the Stages of Development index, dividing countries into different stages of development (Stages 1 to 3) based on the real GDP per capita. The last index is used in this study as a basis for grouping countries into developed countries (Stage 3: GDP per capita > USD 17,000) and less developed countries (Stage 1 and Stage 2: GDP per capita ≤ USD 17,000).

The Economic Freedom of the World Report (Gwartney and Lawson, 2009) provides data on international capital market controls. The related index is scaled 1 to 10 with higher scale indicating more open capital flow policy. It is constructed based on two sub indices: (1) foreign ownership/investment restrictions, and (2) capital controls. The sources of the two sub indices are the WEF's Global Competitiveness Report and the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions. The data on country credit ratings are from various editions of the Institutional Investor (2003-2007). The creditworthiness of each country is evaluated on a scale of 1 to 100 with higher scale indicating higher rating. The rating is constructed based on a compilation of economic, financial, and political indicators assessed by senior economists and risk analysts worldwide.

The data on distances are from Centre d'Etudes Prospectives et d'Informations Internationales (CEPII) Geodesic distances database (2010). CEPII calculates geodesic distances with latitudes and longitudes of the most important cities/agglomerations in terms of population.

1.3 Literature review

(a) Finance and trade literature

The literature indicates that there is a significant relationship between financial development and trade. Beck (2002) provides a theoretical model examining the relationship between financial development and trade, focusing on the role played by financial intermediaries in facilitating high-return manufacturing projects. Using panel data on private credit for 65 countries over a 30-year period, the study finds that financial development strongly affects export volume and trade balance of manufacturing industries (2002, p. 107). In a subsequent paper, Beck (2003) finds that countries in which financial systems are relatively highly developed tend to have higher export shares and trade balances when industries rely more on external finance.

A broad range of literature link international capital flows to trade. Some of the studies use macroeconomic models, commonly based on the adjustments to changes in trade regulations or capital market controls. McKinnon (1993) suggests the importance of capital market controls in determining trade flows. On the other hand, Mundell (1957) examines the effect of trade openness on capital flows. Based on a 2x2 Heckscher-Ohlin factor endowment model, Mundell suggests that barriers to trade encourage international capital flows that if unhindered increase the output of the host country's import-competing sector. Several recent papers on this topic relate trade openness to the stability of capital flows. Cavallo and Frankel (2008) provide empirical evidence that trade openness makes countries less susceptible to sudden stops in capital inflows and thus less susceptible to crises.

Other studies on trade and capital flows utilize microeconomic models, focusing on firm-level problems such as costs and sales. These studies examine specifically the relationship between trade and foreign direct investment (FDI). A model of trade-FDI relationship is known as the "Proximity-Concentration Tradeoff" suggesting that multinational companies choose exporting over FDI when they face higher fixed costs in the host countries than trade costs, but choose FDI over exporting when the trade costs are higher than fixed costs (Smith [1987] and Neary [2002]). It assumes that trade and FDI are substitutes. Neary (2009) argues, however, that the proximity-concentration tradeoff applies to horizontal FDI only.

A vertical FDI model suggests that trade and FDI can be complements if countries differ either in technology or endowments of specific factors (Markunsen [1983] and Neary [1995]). Trade liberalization can encourage FDI if the induced capital flows lead to an increase in production of the host country's exporting sector. Helpman (1984) utilizes a Heckscher-Ohlin model to make a point that when stages of production vary in factor intensities, differences in factor endowments between countries may encourage vertical disintegration by firms.

Based on data of U.S. capital outflows, Ruffin and Rassekh (1986) argue that foreign direct investment and portfolio investment are perfect substitutes. They note that the way multinational corporations finance their operations may be unrelated to the net flow of capital between countries (p. 1126). A recent article on foreign direct investment and portfolio investment is Goldstein and Razin (2005). Using a model of information-based tradeoffs between direct investment and portfolio investment, they compare the expected yields on the two types of foreign investment. Goldstein and Razin point out that developed countries attract more portfolio investment relative to direct investment than less developed countries due to the higher fixed costs associated with direct investment in developed countries. They also note that the high levels of transparency make portfolio investment in developed countries more efficient.

Real exchange rate and its role in determining the profitability of tradables have attracted more attention in the recent economic growth literature. Freund and Pierola (2008) find that a currency undervaluation and a decrease in exchange rate volatility increase exports in developing countries. They also find that exchange rate volatility has less impact on exports in developed countries. It is often suggested that exchange rate volatility reduces trade; however, some empirical studies find that the negative relationship is not robust to some specifications (Clark, Tamirisa, and Wei [2004]; Baxter and Kouparitsas [2006]). Furthermore, Gala (2008, p. 279) points out that "currency misalignment measures are far from consensual" and contributes to the literature by theoretically and empirically connecting real exchange rate levels and economic development. Gala finds that real exchange rate levels could affect economic growth through capital accumulation and technological development.

(b) Gravity literature

Gravity models have been widely used in examining international trade flows. Proposed by Tinbergen (1962), the gravity equation is later developed by several authors who provide the microfoundations (e.g., Anderson [1979], Hepman [1987], Anderson and van Wincoop [2003], and Feenstra [2004]).

The gravity model predicts that trade flows between two countries are inversely related to the physical distance between them and directly related to the multiplicative interaction of each country's size, which is commonly measured by GDP. Anderson (1979) develops a microfoundation for the gravity model using constant elasticity of substitution (CES) expenditure system. It proves algebraically that level of trade is related to the size (income) of countries as well as trade costs. Trade costs include: (1) transport costs as an increasing function of distance, and (2) tariffs. Another application of the gravity equation widely cited in subsequent papers is an empirical work by McCallum (1995) that finds that national borders matter for trade flows. Feenstra (2004) provides a theoretical model explaining border effects, such as transport costs or tariffs, using the CES utility function in a consumer utility maximization problem.

Anderson and van Wincoop (2003) point out the role of prices in the applications of gravity equation for international trade. They find that ignoring prices in the cross-section gravity equation generates omitted variables bias. Using the general equilibrium structure of the model, they calculate the comparative statics of trade barriers and propose a framework that counts for "multilateral price resistance." They use a nonlinear least squares program to estimate the multilateral price resistance terms.

Feenstra (2004) suggests an alternative method to estimate the multilateral price resistance terms in cross section by including country-specific fixed effects, controlling for the effects of exporters by time as well as importers by time. After implementing the country-specific fixed effects, Baier and Bergstrand (2007) find that there is still a large amount of unobserved heterogeneity among country pairs, and thus they add country-pair fixed effects. The country-pair fixed effects capture the effects of free trade agreements (FTA) on trade. Baltagi, Egger, and Pfaffermayr (2003) note that a full interaction effect may be used in examining bilateral trade flows. The full interaction effects consist of country-specific fixed effects, country-pair fixed effects, as well as fixed exporter, importer, and time effects.

The robustness of various variables in a gravity model is examined by Baxter and Koupartisas (2006). They utilize three methods for determining robust relationships: the extreme bounds analysis of Leamer (1983, 1985), the extreme bounds analysis of Sala-i-Martin (1997), and the general to specific approach of Hendry (1995). They find that *common* gravity variables are robust determinant of trade. These variables are distance, common border, cultural distance, and colonial ties. Other variables tested produce different results. Bilateral factor endowment and stage of development are robust determinants of trade when tested using all three methods. Industrial similarity, restrictions to flows of goods and capital, currency union membership, and exchange rate volatility show robustness under only one of the testing methods or when tested with a restricted sample.

1.4 Gravity model

The bilateral trade model is specified with trade flows between country i and country j (T_{ij}) directly related to the multiplicative interaction of each country's size and inversely related to the physical distance between them,

$$T_{ij} = \beta_o\, Y_i^{\beta_1} Y_j^{\beta_2}\, D_{ij}^{\beta_3} \varepsilon_{ij}\ , \qquad (1)$$

where $Y_{i(j)}$ denotes GDP of country i(j) and D_{ij} denotes the distance between country i and j; β_s are unknown parameters, and ε_{ij} is a stochastic error term.

Taking natural logarithms of both sides, the gravity equation can be estimated.

$$\ln T_{ij} = \beta_o + \beta_1 \ln Y_i + \beta_2 \ln Y_j + \beta_3 \ln D_{ij} + \varepsilon_{ij} \qquad (2)$$

Adding Z_{ij} and W_{ij} as continuous variables and dummy variables of interest in the study, the gravity equation is rewritten as follows.

$$\ln T_{ij} = \beta_o + \beta_1 \ln Y_i + \beta_2 \ln Y_j + \beta_3 \ln D_{ij} + \gamma \ln Z_{ij} + \delta W_{ij} + \varepsilon_{ij} \qquad (3)$$

Following Anderson and van Wincoop (2003), a framework that accounts for multilateral (price) resistance terms is built by including $P_i^{1\text{-}\sigma}$ and $P_j^{\,1\text{-}\sigma}$, main country and partner country price indices. The model proposes that the degree of trade is determined not only by the national borders between country i and country j but also by the multilateral resistance from their trading partners in the rest of the world. It also moves the GDP terms from the right to the left side of the gravity equation.

$$\ln\,[T_{ij}/(Y_iY_j)] = \beta_o + \beta_3 \ln D_{ij} + \gamma \ln Z_{ij} + \delta\, W_{ij} - \ln P_i^{\,1\text{-}\sigma} - \ln P_j^{\,1\text{-}\sigma} + \varepsilon_{ij} \qquad (4)$$

where,

$$P_i^{1-\sigma} = \sum_{i=1}^{N} P_i^{\sigma-1} \left(\frac{Y_i}{Y_w}\right) e^{\beta 3\; ln\; Dij\; +\; \gamma ln\; Zij\; +\; \delta\; Wij} \qquad (5)$$

with $i = 1 \ldots N$ equilibrium conditions; Y_w denotes world GDP (constant across countries); and σ is the elasticity of substitution between countries.

This paper follows the measure of bilateral trade by Helpman (1987) and Feenstra (2004). In a panel setting, they define bilateral trade between country i and country j in period t, T_{ijt}, as:

T_{ijt}= ln(X_{ijt} + X_{jit}), where X_{ijt} denotes exports from country i to country j in period t and X_{jit} denotes the reverse.

The conceptual model is as follows.

$$Trade = \mathrm{f}(Gravity, External\ Funds, International\ Finance, Other) \qquad (6)$$

where,

Trade is the level of trade between pair countries relative to their level of GDP.

Gravity is a group of variables commonly used in gravity equations.

External Funds is a group of variables indicating access to external funds.

International Finance is a group of international financial indicators.

Other is a group of other variables controlling for fixed effects.

[Figure 1.1 here]

The variables of interests of this study are the financial development variables, represented by Z_{ij} in (4). These variables comprise the External Funds variables (access to loans and access to the local equity market) and the International Finance variables (country credit ratings, international capital market controls, and the real exchange rates). Dummy variables commonly used in gravity equations are represented by W_{ij} in (4). These variables are land common border and stage of development.

The present study uses a fixed effects method to estimate the gravity equation, controlling for country-specific (by time) fixed effects and country-pair fixed effects. The country-specific fixed effects count for the multilateral resistance terms as suggested by Anderson and van

Wincoop (2003) and Feenstra (2004), whereas the country-pair fixed effects capture the effects of free trade agreements (FTA) as suggested by Baier and Bergstrand (2007).

The gravity equation for trade and financial development is specified as follows.

$$TRA_{ijt} = \beta_o + \beta_1 DIS_{ij} + \beta_2 LCB_{ij} + \beta_3 DEV_{ijt} + \beta_4 LNA_{ijt} + \beta_5 EQA_{ijt} + \beta_6 CCR_{ijt} + \beta_7 CAP_{ijt} + \beta_8 RER_{ijt} + \gamma_{it} + \delta_{jt} + \alpha_{ij} + \varepsilon_{ijt} \quad (7)$$

where i denotes main countries, j denotes partner countries, t denotes time; and the variables are described as follows.

TRA_{ijt} = log of trade between i and j relative to the product of their levels of GDP.

DIS_{ij} = log of the distance between i and j.

LCB_{ij} = a dummy variable which is unity if i and j share a common land border.

DEV_{ijt} = a dummy variable which is unity if i and j are in the same stage of development.

LNA_{ijt} = log of the product of the scores of access to loans in i and j.

EQA_{ijt} = log of the product of the scores of access to local equity market in i and j.

CCR_{ijt} = the difference between the log of the maximum of the country credit ratings of i and j and the log of the minimum of the country credit ratings of i and j.

CAP_{ijt} = log of the product of the scores of capital controls in i and j.

RER_{ijt} = the difference between the log of the maximum of the real effective exchange rates of i and j and the log of the minimum of the real effective exchange rates of *i* and *j*.[3]

γ_{it} = interactions between country *i* and year *t* dummies

δ_{jt} = interactions between country *j* and year *t* dummies

α_{ij} = interactions between country *i* and country *j* dummies

ε_{ijt} = other influences on bilateral trade

The financial development variables are measured with bilateral interactions similar to those of the common gravity variables. Two forms of bilateral interaction are constructed for the financial development variables following the methods used by Baxter and Kouparitsas (2006) in measuring factor intensity.

The first measure is the log of the product of the levels of 'factor' in the two countries,

$$\text{Bilateral interaction 1: } \ln(F_{it} * F_{jt}) \qquad (8)$$

where $F_{it(jt)}$ stands for the factor in country *i*(*j*) in period *t*. This measure of 'scale' can be interpreted as: "the higher the measure, the higher the levels of factor in the two countries or in either of the two countries." It also suggests that "the higher the measure, the more equal the levels of factor in the two countries." This interaction term is applied to the measures of access to loans (*LNA*), access to equity market (*EQA*), and international capital market controls (*CAP*).

The second measure is the difference between the log of the maximum and the log of the minimum levels of factor in the two countries, or

[3] Real exchange rate is defined as: $RER = e \times P_{FC}/P_{DC}$, where e is nominal exchange rate (DC/FC) and P_{FC}/P_{DC} is the ratio of price levels in the two countries. DC denotes domestic currency, and FC denotes foreign currency. However, the data used are real effective exchange rates, the averages of bilateral real exchange rates between the country and each of its trading partners, weighted by the respective trade shares of each partner. The IMF's real effective exchange rates index (based on relative consumer prices) is used to allow for comparison with a broad range of partner countries.

$$\text{Bilateral interaction 2: } \ln[\max(F_{it}, F_{jt})/\min(F_{it}, F_{jt})] \quad (9)$$

This measure can be interpreted as follows: "the larger the measure, the larger the difference in the levels of factors in the two countries." This interaction term is applied to the measures of country credit ratings (*CCR*) and real exchange rates (*RER*).

Equation (8) is based on a similar method of measuring the scale of GDP for pair countries, whereas equation (9) is often used in gravity literature to measure the differences in GDP per capita between pair countries. Baxter and Kouparitsas (2006) also use both equations as the measures of human capital.

[Table 1.2 here]

1.5 Empirical results

The gravity model described in Section 1.4, based on Anderson and van Wincoop (2003) and Baier and Bergstrand (2007), provides estimation results as presented in Table 1.4(a). As comparisons, Table 1.4(b) and Table 1.4(c) list the regression results obtained from using only country-specific by time fixed effects (Anderson and van Wincoop, 2003) and Ordinary Least Squares (OLS), respectively.

Ignoring country-pair fixed effects, the estimates in Table 1.4(b) generate lower values of adjusted *R*-squared compared to the estimations of the selected model (Table 1.4[a]). The estimates in Table 1.4(b) are biased due to the omission of the effects of FTA when in fact the countries are affected by at least two major FTAs (European Union and NAFTA).

The OLS estimates (Table 1.4[c]) show the lowest values of adjusted *R*-squared among the three models. The regression for Dataset 1 shows that stage of development variable has a significant negative relationship with bilateral trade at 1% level of significance. This result suggests that the dataset should be partitioned based on partner country's stage of development: developed countries (Dataset 2) and less developed countries (Dataset 3). In addition, a Chow test is conducted, and the result rejects the null hypothesis of no differences between the two groups at 1% level of significance.[4]

Table 1.5 lists the estimation results obtained from using financial development variables without bilateral interactions (e.g., using LNA_{it} and LNA_{jt} instead of LNA_{ijt}). It implies that the coefficient estimate for a finance variable for country *i*(*j*) is generated by holding constant the same variable for country *j*(*i*) and other variables in the model. This model seems to work for data on less developed countries, but not on developed countries.

[Tables 1.3, 1.4, and 1.5 here]

The following paragraphs discuss separately and in details the regression results for developed and less developed countries.

(a) Developed (partner) countries

A developed (partner) country is a trading partner that shares the same stage of development as the three main countries, represented by a dummy variable that is in unity. Panel data of developed countries, named Dataset 2, comprise 318 observations (Table 1.4[a]). The adjusted *R*-squared shows that 95.8% of the variation in bilateral trade is explained by the model. An *F*-test is conducted for each group of variables and it shows that the Gravity variables (*DIS* and *LCB*) jointly contribute to bilateral trade at 1% level of significance. The External Funds

[4]Chow test, *F*-value = 3.27.

variables (*LNA* and *EQA* jointly) and the International Finance variables (*CCR*, *CAP*, and *RER* jointly) also significantly contribute to bilateral trade at 1% level. [5]

Distance between pair countries has a negative relationship with bilateral trade at 1% level of significance. Main countries trade more with developed countries located closer to them. Land common border indicates a positive relationship with trade between developed countries, also significant at 1% level.

Access to loans has a positive relationship with bilateral trade at 5% level of significance, suggesting that trade between countries increases when it is easier for businesses to obtain loans. Access to equity shows a negative relationship with bilateral trade at 1% level of significance, indicating that trade between countries decreases when it is easier for businesses to raise capital through the local equity markets. Since the pair countries in this data set are both developed countries, this result seems to be intuitively correct. The rationale for this result may relate to the degree of efficiency of the equity markets. In developed countries, the equity markets have high levels of transparency and thus more efficient than those in less developed countries (Goldstein and Razin, 2005). Easy access to highly efficient equity markets may encourage multinational companies to substitute trade with investment in local suppliers. Thus, easier access to equity market lowers trade.

The country credit ratings variable indicates a strong negative relationship with bilateral trade at 1% level of significance. The larger the difference in credit ratings of the two countries, the less trade occurs between them. The capital controls variable also shows a strong negative relationship with bilateral trade at 1% level of significance. It suggests that trade between two

[5] Gravity, *F*-value=83.58; External Funds, *F*-value=7.64; International Finance, *F*-value=8.50.

countries decreases when the countries relax their international capital market controls. This result indicates a substitution relationship between international trade and foreign investment in local companies, consistent with the horizontal FDI model proposed by Smith (1987) and Neary (2002). The real exchange rate variable is positively related to bilateral trade at 1% level of significance, indicating the larger the difference in real exchange rates of pair countries, the more trade occurs between them.

(b) Less developed (partner) countries

A less developed (partner) country is a trading partner that is in a lower stage of development than the three main countries. The panel data of less developed partner countries, named Dataset 3, comprise 372 observations (Table 1.4). The adjusted R-squared shows that 93.6% of the variation in bilateral trade is explained by the model. The F-test for each group of variables shows that Gravity variables (*DIS* and *LCB*) jointly contribute to bilateral trade at 1% level of significance. The External Funds variables (*LNA* and *EQA* jointly) and International Finance variables (*CCR*, *CAP*, and *RER* jointly) are also significant at 1% level.[6]

As expected, distance between pair countries shows a strong negative relationship with trade at 1% level of significance. Land common border, however, does not significantly contribute to bilateral trade when estimated using this data set.

Access to loans has a positive relationship with bilateral trade at 1% level of significance. The degree of bilateral trade contributed by access to loans is higher when estimated using this data set compared to using Dataset 2. Access to equity also shows a positive relationship with

[6] Gravity, F-value=86.76; External Funds, F-value=16.71; International Finance, F-value=48.87.

bilateral trade, significant at 5% level, indicating that trade increases when it is easier for businesses in the two countries to raise capital through the local equity market.

The country credit ratings variable is negatively related to bilateral trade at 1% level of significance, whereas the capital controls variable is positively related to bilateral trade at 1% level of significance. In contrast to the negative result from Dataset 2, the positive coefficient suggests that trade between main countries and less developed partner countries increases when these countries relax their international capital market controls. This result is consistent with the vertical FDI model developed by Markunsen (1983) and Neary (1995). The real exchange rates variable shows a negative relationship with bilateral trade at 5% level of significance, indicating that the larger the difference in real exchange rates of pair countries, the less trade occurs between them.

The regression without pair-country interaction terms for financial development variables reveals some interesting results (Table 1.5). An increase in real effective exchange rates in less developed countries seems to increase trade with main countries at 1% level of significance, holding constant real effective exchange rates in main countries. This result is consistent with the literature (Freund and Pierola, 2008). A real depreciation of currency in the main countries, however, decreases trade with less developed countries at 10% level of significance. Other coefficient estimates appear to be consistent with the estimation results from the model with interaction terms (Table 1.4a), except for access to loans variables (LNA_{it} and LNA_{jt}) that show insignificant results.

1.6 Conclusion

This study investigates the relationship between financial development and trade using a gravity model with fixed effects based on Anderson and van Wincoop (2003) and Baier and Bergstrand (2007). The panel data of bilateral trade between the three main countries (United States, Japan, and Germany) and 47 partner countries covers the period 2003 to 2007. The study finds differences when grouping the partner countries into developed and less developed countries affecting the degree of bilateral trade. Consistent with the literature, physical distance is negatively related to trade between countries.

In all cases, access to loans has a significant positive relationship with bilateral trade. Easy access to loans in less developed countries appears to contribute more to the bilateral trade than in developed countries. This result is expected since financial systems in less developed countries are heavily bank based.

Access to equity shows a significant positive relationship with bilateral trade for less developed countries. Equity market development, as well as trade, is often promoted in less developed countries since financial systems in these countries are usually bank based rather than market based. In contrast, trade between main countries and other developed countries is lower with higher access to equity. A possible explanation is that the equity markets in developed countries have high levels of transparency and are more efficient than the equity markets in less developed countries. Easy access to highly efficient equity markets in developed countries may encourage multinational companies to substitute trade with investment in local suppliers, lowering trade.

All three international financial indicators (country credit ratings, capital controls, real exchange rates) have significant relationships with bilateral trade. The three variables appear to affect trade between main countries and less developed countries to a lesser degree. In all cases, main countries tend to trade more with partner countries with higher credit ratings.

Trade between main countries and other developed countries is lower when the countries relax their international capital market controls. This result suggests substitution between trade and foreign investment in local companies that serve the local markets, consistent with horizontal FDI model. On the other hand, trade between main countries and less developed countries is higher when the countries relax their capital controls, consistent with vertical FDI model.

The real exchange rate is positively related to bilateral trade when main countries trade with other developed countries. The positive relationship indicates that a larger difference in real exchange rates between two developed countries implies more trade. The rationale may relate to a shift in production location of multinational firms within developed countries to take advantage of misaligned currency. The opposite is the case for trade with less developed countries. This result, however, needs to be interpreted with caution and merits further study.

In conclusion, this paper finds that there is indeed a significant relationship between financial development and trade. The results suggest implications for policies regarding access to external funds for businesses as well as capital controls and exchange rates in promoting trade and development, with some differences depending on whether the country is a developed or less developed country.

CHAPTER 2

Commodity Prices and the Exchange Rate

2.1 Introduction

The global pricing of primary commodities is determined in the world markets connecting the supply of and demand for the goods. Consequently, time series data on these prices are expected to exhibit stationary or mean reverting process, reflecting the dynamic stability of market equilibria. However, empirical unit root tests of international commodity prices generally exhibit highly persistent or even nonstationary process for both individual and aggregate commodities. This phenomenon provides opportunity to explore what seems to generate the inconsistency between theory and empirical evidence. The present study investigates what factors determine commodity prices and provides a rationale for how these factors generate the inconsistency between theory and empirical evidence.

Previous studies have observed this inconsistency between economic theory and empirical evidence suggested by the unit root test results on commodity prices. Wang and Tomek (2007) note that commodity prices should be level stationary according to price theory. Kellard and Wohar (2006) point out that commodity prices should be trend stationary to be consistent with the Prebisch-Singer hypothesis. They further report some evidence of nonlinear stationarity of commodity prices. The present paper uses a different approach in that it supports the evidence of nonstationarity of commodity prices and attempts to filter its factor. The proposed inference is: "factoring out the nonstationary effect will generate the filtered commodity prices that are consistent with economic theory."

An important factor that is likely to explain a nonstationary effect of international commodity prices is the U.S. nominal exchange rate. The U.S. dollar is used in the global pricing of most internationally traded commodities. It is widely accepted that the U.S. nominal exchange rate, e.g., relative to the Euro, Japanese Yen, or some trade-weighted index of currencies, is nonstationary. Its relationship with goods priced in U.S. dollars can be explained as follows. A depreciation of the U.S. dollar should drive an increase in the price of the goods priced in U.S. dollars to maintain the same world price. Another possible explanation is that a depreciation of the U.S. dollar results in lower commodity prices in terms of the foreign currency; and as the foreign demands for the commodities increase, their prices rise.

Based on the above rationales, commodity prices should exhibit a dynamic behavior that closely mirrors the dynamics of the U.S. exchange rate and thus reflect its nonstationarity, which is *common* to all commodity prices. Furthermore, this inference holds for not only nominal commodity prices but also relative commodity prices, which are prices deflated by the U.S. Consumer Price Index (CPI). This similar dynamics of relative prices and nominal prices is expected since aggregate price indices such as the CPI exhibit less fluctuation than international commodity prices.

This study begins with a factor analysis on a panel of 51 international commodity prices for the period January 1980 to December 2009 by performing the PANIC (Panel Analysis of Nonstationarity in Idiosyncratic and Common Components) method, which is a second-generation panel unit root test recently developed by Bai and Ng (2004).[7] An out-of-sample forecasting analysis is then conducted to further investigate the link between the U.S. nominal exchange rate and commodity prices. The present paper examines the predictive power of international commodity prices for movements in the U.S. exchange rate.

[7] PANIC method (Bai and Ng 2004) is chosen over other second-generation panel unit root tests such as Phillips and Sul (2004), Moon and Peron (2004), and Pesaran (2007) because these methods assume stationary common factors and thus do not apply to commodity prices based on the inference regarding the role of the U.S. nominal exchange rate.

The PANIC test results identify two common factors from both the nominal and relative commodity prices. The test results indicate that the first (most important) common factor is nonstationary, whereas the second common factor and the defactored (filtered) idiosyncratic components are both stationary. Graphically, the first common factor reflects the inverse of the U.S. nominal exchange rate. Out-of-sample forecasts using a simple model with the two common factors outperforms a benchmark random walk model. This forecasting result further supports the inference that the first common factor reflects the effect of the U.S. nominal exchange rate on commodity prices. The stationarity of the second common factor and the idiosyncratic components supports the existence of dynamic stability of market equilibria. Overall, these results provide a rationale for the previously identified inconsistency between economic theory and empirical evidence.

This paper proceeds as follows. Section 2 describes the data and their sources. Section 3 provides literature review. Section 4 and section 5 discuss the econometric techniques and the empirical results. Section 6 concludes the paper.

2.2 Data

The data set is constructed based on monthly data of U.S. nominal exchange rates and commodity prices for the period January 1980 to December 2009. The U.S. nominal exchange rate data are the trade-weighted exchange rate index for the U.S. dollar against a basket of major currencies including the Euro, Canada, Japan, the United Kingdom, Switzerland, Australia, and Sweden. The data for 51 commodity prices are from the IMF Primary Commodity Prices database available on the International Monetary Fund Website, except for the natural gas price from the U.S. Energy Information Administration Website.

[Table 2.1 here]

2.3 Literature review

(a) Commodity prices and the exchange rate

There have been attempts to explain the relationship between commodity prices and exchange rates, including those seeking to produce forecasts of commodity price movements. Existing economic models are rarely successful in predicting the exchange rate itself. Meese and Rogoff (1983) point out that nominal exchange rate models routinely fail to outperform the random walk model in predicting out-of-sample in the floating regime.

The notion that commodity prices are perfectly arbitraged is also in question. One related literature is empirical work on the validity of the law of one price (LOP) in commodity markets. Since seminal work of Isard (1977), some find evidence against the LOP (e.g., Ardeni [1989], Engel and Rogers [2001], Parsley and Wei [2001], and Goldberg and Verboven [2005]), while others find evidence in favor of the LOP (e.g., Goodwin [1992], Michael *et al.*[1994], Obsfeld and Taylor [1997], Lo and Zivot [2001], and Sarno *et al.*[2004]).

Several studies observe the inconsistency between economic theory and empirical evidence suggested by the unit root test results on commodity prices. Wang and Tomek (2007) note that price theory suggests that commodity prices should be level stationary. Kellard and Wohar (2006) point out that Prebisch-Singer hypothesis implies that commodity prices should be trend stationary. They recognize the weakness of conventional unit root tests, due to low power of the tests, and provide some evidence of nonlinear stationarity of commodity prices.

A growing body of literature investigates the information contents in commodity prices and other macroeconomic variables. Chen, Rogoff, and Rossi (2008) find substantial evidence of the predictive power of commodity-currency exchange rates for commodity prices, both in-sample and out-of-sample, but no evidence going in reverse direction. Using the exchange rate approach of the preceding researchers as well as a broader approach that includes exchange rates and other macrovariables, Groen and Presenti (2010) find that neither of both approaches has strong predictive power of exchange rates for commodity prices. They note, however, that both approaches occasionally outperform simple benchmark models. Gospodinov and Ng (2010) examine the link between commodity prices and inflation. They provide strong evidence that information contents in commodity prices have predictive power for inflation.

(b) Panel unit root tests

The application of panel unit root tests provides substantial advantages over the use of univariate unit root tests that in practice lack power by the short period of macroeconomic time series. Panel unit root tests increase the power of unit root tests by pooling information across units. Levin, Lin, and Chu (2002) provide evidence of the increase in power generated by panel unit root tests. Madala and Wu (1999) and Im, Pesaran, and Shin (2003) also propose the use of panel unit root tests instead of univariate unit root tests. The first generation unit root tests, however, assume independent units. The drawback of imposing this assumption is that the test results suffer from serious size distortions if the panel data are in fact cross section dependence.

Pesaran (2004) proposes a test for cross section dependence based on averages of pair-wise correlation coefficients of the OLS residuals from the individual regressions in the panel. The test is based on an assumption of no a priori ordering of the cross section units. In a subsequent paper, Pesaran (2007) implements a cross-sectionally augmented Dickey-Fuller

(CADF) test. The CADF augments the standard ADF tests with the cross-section averages of lagged levels and first-differences of the individual series. Pesaran shows that panel unit root tests that do not account for cross section dependence are seriously biased if the degree of dependence is sufficiently large.

The second generation unit root tests allow for cross section dependence among units and thus have better size property than the first generation tests. They suggest the presence of *common* effects, which are assumed to be stationary by Philips and Sul (2003), Moon and Perron (2004) and Pesaran (2007). The PANIC method of Bai and Ng (2004) use no a priori assumption on stationary or integrated processes of the common effects. It utilizes the factor structure of large dimensional panels to examine the nature of nonstationarity in the data. PANIC tests the unobserved components of the data instead of the observed series. Bai and Ng show that, if a factor structure exists in a panel data set, testing the presence of unit roots in the common factors and the idiosyncratic errors separately should be more effective than testing for individual unit roots in the observed series.

2.4 Econometric techniques

This section illustrates the econometric techniques used in this paper. First, augmented Dickey-Fuller (ADF) tests for univariate unit roots are applied to each unit in the data set. The ADF test results provide an assessment of the time series properties and the number of integrated time series in the cross section. The cross section dependence test by Pesaran (2004) is applied to the residuals of the ADF regressions to determine whether there is cross section dependence among commodity prices. Since the present study finds evidence of cross section dependence, it justifies the use of the second-generation panel unit root test of Bai and Ng (2004).

Test for cross section dependence (*CD*) by Pesaran (2004) is based on the averages of pair-wise correlation coefficients of the OLS residuals from the individual regressions in the panel,

$$CD = \left(\frac{2T}{N(N-1)}\right)^{1/2} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{i,j}\right) \xrightarrow{d} N(0,1) \qquad (1)$$

where $\hat{\rho}_{i,j}$ is the pair-wise correlation coefficients from the residuals of the ADF regressions.

The above statistic has exactly mean zero for fixed values of *T* and *N*, under a wide class of panel data models. A rejection of the null of no cross-section dependence implies that the utilization of second-generation panel tests is appropriate and that the first-generation panel tests should not be used. Pesaran (2004) demonstrates that the test is valid under fairly general conditions even when *T* is small and *N* is large.

The number of common factors is determined using $PC(r)$ and $IC(r)$ criteria of Bai and Ng (2002). The PANIC method of Bai and Ng (2004) is implemented to analyze the nonstationarity characteristics of commodity prices. A conventional method for out-of-sample forecast accuracy proposed by Diebold and Mariano (1995) and West (1996) is used to further examine the link between commodity prices and the U.S. nominal exchange rate.

(a) The PANIC method

The PANIC method of Bai and Ng (2004) is described as follows. Let $p_{i,t}$ be the natural logarithm price of a good i at time t that obeys the following stochastic process[8]

$$p_{i,t} = c_i + \lambda_i' \mathbf{f}_t + e_{i,t} \qquad (2)$$

[8] All regularity conditions in Bai and Ng (2004, pp. 1130-1131) are assumed to be satisfied.

$$(1 - L)\,\mathbf{f}_t = \mathbf{A}\,(L)\,\mathbf{u}_t$$

$$(1 - \rho_i L) e_{i,t} = B_i(L)\varepsilon_{i,t}$$

where c_i is a fixed effect intercept, $\mathbf{f}_t = [f_1 \dots f_r]'$ is a $r \times 1$ vector of (latent) common factors, $\lambda_t = [\lambda_{i,1} \dots \lambda_{i,r}]'$ denotes a $r \times 1$ vector of factor loadings for good i, and $e_{i,t}$ is the idiosyncratic error term. $\mathbf{A}(L)$ and $B_i(L)$ are lag(L) polynomials. $\mathbf{u}_t$, $\varepsilon_{i,t}$, and λ_i are mutually independent. An important point is that a factor model with *N* variables has *N* idiosyncratic components but a small number of common factors ($r << N$).

A factor structure determines that the series have nonstationary process if one or more of the common factors are nonstationary, or the idiosyncratic error is nonstationary, or both (Bai and Ng [2004, p. 1128]). Estimations are performed by the method of principal components. When $e_{i,t}$ is stationary, the principal component estimators for $\mathbf{f}_t$ and λ_i are consistent regardless of the order of $\mathbf{f}_t$. When $e_{i,t}$ is integrated, however, the estimator is inconsistent because a regression of $p_{i,t}$ on $\mathbf{f}_t$ is spurious. Implementing the method of principal components to the first-differenced data, PANIC tackles this problem as follows.

Rewrite (1) as a model with differenced variables,

$$\Delta p_{i,t} = c_i + \lambda_i' \Delta \mathbf{f}_t + \Delta e_{i,t} \tag{3}$$

for $t = 2, \dots, T$. Let $\Delta \mathbf{p}_{\mathrm{i}} = [\Delta p_{\mathrm{i},1} \dots \Delta p_T]'$ and $\Delta \mathbf{p} = [\Delta \mathbf{p}_1 \dots \Delta \mathbf{p}_N]$. After proper normalization[9], the method of principal components for $\Delta \mathbf{p} \Delta \mathbf{p}'$ yields estimated factors $\Delta \hat{\mathbf{f}}_t$, the

[9] This is because the principal components method is not scale invariant.

associated factor loadings $\hat{\lambda}_i$, and the residuals $\Delta\hat{e}_{i,t} = \Delta p_{i,1} - \lambda_i'\Delta\mathbf{f}_t$. Re-integrating these, we obtain

$$\hat{e}_{i,t} = \sum_{s=2}^{t} \Delta\hat{e}_{i,s} \tag{4}$$

for $i = 1, \dots, N$ and

$$\hat{\mathbf{f}}_t = \sum_{s=2}^{t} \Delta\hat{\mathbf{f}}_s \tag{5}$$

Theorem 1 of Bai and Ng (2004, p.1134) shows that $\hat{e}_{i,t}$ and $\hat{\mathbf{f}}_t$ can be tested as if they are observable. Specifically, the ADF test with no deterministic terms can be applied to each $\hat{e}_{i,t}$ and the ADF test with an intercept can be used for $\hat{\mathbf{f}}_t$. When there are more than two nonstationary factors, cointegration-type tests can be used to determine the rank of $\mathbf{A}(1)$ in (2). Bai and Ng (2004) proposed a panel unit root test for idiosyncratic terms as follows.

$$P_{\hat{e}} = \frac{-2\sum_{i=1}^{N} \ln p_{\widehat{e_i}} - 2N}{2N^{1/2}} \xrightarrow{d} N(0,1) \tag{6}$$

where $p_{\widehat{e_i}}$ is the *p*-value from the ADF test for $\hat{e}_{i,t}$. Pooling these *p*-values across units tolerates the presence of as much heterogeneity across units as possible.

The estimates for common factors $(\widehat{\mathbf{f}}_t)$, factor loadings $(\hat{\lambda}_i)$, and idiosyncratic components $(\hat{e}_{i,t})$ are obtained by applying the method of principal components. The importance of common factors for dynamics of the commodity prices relative to idiosyncratic components is evaluated by

$$rv_i^k = \frac{\sigma\left(\hat{\lambda}_i \hat{f}_t^k\right)}{\sigma\left(\hat{e}_{i,t}\right)}, k = 1, \dots, r \tag{7}$$

where $\sigma(\cdot)$ denotes the standard deviation.

(b) Diebold-Mariano-West Test Statistics

Out-of-sample forecasts of the U.S. nominal exchange rate are performed using a model based on the estimated common factors with the random walk model as a benchmark. A conventional method developed by Diebold and Mariano (1995) and West (1996) is used to evaluate the out-of-sample forecast accuracy of these models.

Let s_t denote the natural logarithm U.S. nominal exchange rate. The random walk model of s_t implies,

$$s^R_{t+k|t} \; = \; s_t, \tag{8}$$

where $s^R_{t+k|t}$ is the *k*-step ahead forecast by the random walk model given information set at time *t*. The competing factor model is based on the following least squares regression.

$$\mathbf{\Delta} s_{t+k} = \; c + \; \beta' \mathbf{\Delta f}_t + \; u_t \tag{9}$$

Given the least squares coefficient estimate, the *k*-step ahead forecast by the factor model $s^F_{t+k|t}$ is

$$s^F_{t+k|t} \; = \; \textstyle\sum_{s=1}^{k} \widehat{\mathbf{\Delta} s_{t+s}} + s_t, \tag{10}$$

where $\widehat{\mathbf{\Delta} s_{t+s}}$ is the fitted value from (7) and s_t is the actual data at time *t*.

The forecast errors from the two models are

$$\varepsilon^R_{t+k|t} = \; s_{t+k} - \; s^R_{t+k|t}, \qquad \varepsilon^F_{t+k|t} = \; s_{t+k} - \; s^F_{t+k|t}$$

The loss differentials are defined as

$$d_t = L\big(\varepsilon^R_{t+k|t}\big) - L(\varepsilon^F_{t+k|t})$$

where $L(\varepsilon^{j}_{t+k|t})$, $j = R, F$ is a loss function.[10] To test the null of equal predictive accuracy, H_0: $Ed_t = 0$, the Diebold-Mariano-West statistic (*DMW*) is defined as

$$DMW = \frac{\bar{d}}{\sqrt{\widehat{Avar}(\bar{d})}} \tag{11}$$

where $\bar{d}$ is the sample mean loss differential,

$$\bar{d} = \frac{1}{T - T_0} \sum_{t=T_o+1}^{T} d_t$$

$\widehat{Avar}(\bar{d})$ is the asymptotic variance of $\bar{d}$,

$$\widehat{Avar}(\bar{d}) = \frac{1}{T - T_0} \sum_{j=-q}^{q} k(j,q)\hat{\Gamma}_j \ ,$$

and $k(\cdot)$ denotes a kernel function where $k(\cdot) = 0, j > q$, and $\hat{\Gamma}_j$ is j^{th} autocovariance function estimate.[11] It is known that the DMW statistic is severely undersized with asymptotic critical values when competing models are nested, which is the case here. Therefore, the critical values of McCracken (2007) are used to avoid this size distortion problem.

2.5 Empirical results

As a preliminary analysis, the ADF test for individual commodity prices are performed (see Table 2.2). The test rejects the null of nonstationarity for only 13 and 14 out of 51

[10] We use the conventional squared error loss function, $(\varepsilon^{j}_{t} + k|t)^2, \ j = R, F$.

[11] Following Andrews and Monahan (1992), the quadratic spectral kernel with automatic bandwidth selection is used in the analysis.

commodity prices at the 5% significance level. Since the ADF test is known to suffer from low power in small samples, this test result cannot be taken as an evidence for overall nonstationarity of commodity prices. The cross-section dependence (CD) test of Pesaran (2004) rejects the null of no cross-section dependence at any significance level (see Table 2.2), which justifies the use of second-generation unit root tests and suggests that using the first-generation panel tests is not appropriate.

[Table 2.2 here]

The next step is to implement PANIC for the commodity prices. Using $PC(r)$ and $IC(r)$ criteria of Bai and Ng (2002), the number of common factors is determined. Figure 2.1 indicates that all criteria except $PC_3(r)$ choose two factors ($r = 2$).

[Figures 2.1and 2.2 here]

Figures 2.3 and 2.4 show that dynamics of individual commodity prices are substantially influenced by the first common factor. For many commodity prices, rv_i^k is greater than one, suggesting that the first common factor is more important than idiosyncratic components for those prices. The second common factor also plays an important role for some commodities such as crude oil prices. Similar evidence can be found in factor loading estimates (Figures 2.5 and 2.6).

[Figures 2.3 through 2.6 here]

The PANIC unit root test results are reported in Table 2.3. The ADF test fails to reject null of nonstationarity for the first factor (f_t^1), but rejects the null for the second factor (f_t^2) at 5% level of significance. The cointegration tests are not implemented in this case since there is only one nonstationary factor among two common factors, $rank\left(\mathbf{A}(1)\right) = 1$. For the defactored (filtered) idiosyncratic components, the ADF tests reject the null for 30 and 29 out of 51 nominal and relative commodity prices, respectively. The panel unit root test by (6) rejects the null hypothesis at any significance level. These results provide strong evidence that there is a nonstationary common factor that drives persistent movement of commodity prices.

[Table 2.3 here]

Although there is no obvious way of identifying the source of this nonstationarity (since the factors are latent variables), it can be observed that the estimated first common factor mirrors the image of the U.S. nominal exchange rate. Figure 2.3 shows that the exchange rate exhibits two large swings peaked in 1985 and in 2002, whereas the first common factor estimate shows similar patterns but in opposite directions. This phenomenon may make sense since it is previously acknowledged that most commodities are priced in dollar terms. A depreciation of U.S. dollar relative to other major currencies may raise commodity prices given the world price, and the reverse holds. The second common factor shows stable fluctuations which may indicate the stationary characteristics.

The two common factors considerably affect oil prices (Figures 2.3 through 2.6). The three crude oil prices (Brent, Dubai, and WTI) in Figure 2.8 show some interesting dynamics illustrating the role of the common factors in these prices. Original oil prices are plotted in panel (a), and the defactored oil prices are in panel (b). Panel (a) clearly shows extremely persistent movements of oil prices. Defactored oil prices, however, exhibit much less persistent dynamics.

The nonstationarity of nominal exchange rates seems to be widely accepted in the economic profession. If this is the case, and if commodity prices are largely governed by a single nonstationary common factor, it might be the case that such nonstationarity is due to the U.S. nominal exchange rate. The remaining factors and/or idiosyncratic components may reflect changes in world demand and supply conditions, and their movements around the long-run equilibrium may be consistent with price theories.

Lastly, the out-of-sample forecast results reported in Tables 6 and 7 provide additional evidence to support the possibility that the first common factor reflects the effect of exchange rate on commodity prices. Forecasting is conducted recursively by sequentially adding one additional observation from 180 initial observations toward 360 total observations for k = 1, 2, 3, 4. First, the ratios of the root mean square prediction error (RMSPE) of the random walk model to the factor model were greater than one for all k, suggesting that the factor model outperforms the benchmark random walk model. Second, the *DMW* statistics with McCracken's (2007) critical values reject the null of equal predictability for k = 1, 4 at the 5% significance level and for k = 3 at the 10% level.

2.6 Conclusion

This paper examines common factors affecting the prices of 51 highly tradable world commodities. It recognizes an inconsistency between the implications of economic theory concerning the dynamic behavior of commodity prices and the implications of corresponding empirical tests as reported by previous studies (e.g., Wang and Tomek [2007]; Kellard and Wohar [2006]). Dynamic stability of market equilibria implies that time series data on commodity prices should exhibit a stationary or mean reverting process, but unit root tests on these prices generally show evidence of nonstationarity. This paper investigates this inconsistency between economic theory and empirical evidence by first identifying common factors driving the dynamics of highly tradable commodity prices.

Utilizing the PANIC method of Bai and Ng (2004), the study finds that the factor structure of commodity prices is well described by two common factors. The first common factor explains the largest proportion of the variation in the panel of commodity prices. It is nonstationary and closely related to the U.S. nominal exchange rate, as suggested by theoretical, graphical, and out-of-sample forecasting evidence. The second common factor and the defactored (filtered) idiosyncratic components are stationary.

A simple model constructed with the two common factors significantly outperforms a random walk in forecasting the exchange rate. This is an important result because the random walk model consistently outperforms economic models for forecasting the exchange rate since it was reported by Meese and Rogoff (1983). Furthermore, the forecasting result indicates that the first common factor and exchange rates share information content. It suggests that factors that have a predictive power for the exchange rate will have a correspondingly predictive power for commodity prices.

The stationarity of the second common factor and the idiosyncratic components of each series are consistent with equilibrium price dynamics, showing a mean reverting process. When the effects the first common factor, which has been identified to be closely related to the U.S. exchange rate, are filtered out of the panel of commodity prices, the remaining factors affecting commodity prices exhibit the type of dynamic behavior consistent with price theory. An obvious illustration is demonstrated by the three crude oil prices (Brent, Dubai, and WTI), which show extremely persistent (possibly nonstationary) movements of their original prices, but much less persistent dynamics (possibly stationary) movements of their defactored (filtered) prices.

Overall, the results of this study support the proposed rationale for the inconsistency between economic theory and empirical evidence on international commodity prices.

CHAPTER 3

Japanese and Korean Automobile Exports to the United States:

Shipping the Good Cars Out?

3.1 Introduction

This paper analyzes Japanese and Korean automobile exports to the United States based on monthly data for the period 1995 to 2008 to examine consistency with the Alchian-Allen theorem (1964). Also called "shipping the good apples out," the theorem suggests that a per-unit charge applied to two similar goods will lower the relative price and increase the relative consumption of the higher quality good (e.g., Hummels and Skiba [2002]; Bauman [2004]).

Japan has historically been the largest automobile exporter to the United States with a share of about 50% to 70% of U.S. automobile imports during the period of observations, whereas Korea has become one of the major automobile exporters to the United States with a share of about 14% of U.S. imports in 2008.[12] One of the implications of the Alchian-Allen theorem may be related to the choices of production location as major Japanese and (more recently) Korean automobile manufacturers own production plants in the United States.

The present study utilizes the dynamic OLS (DOLS) method of Stock and Watson (1993) to estimate car imports from Japan and Korea with car price, shipping costs, and the U.S. dollar exchange rate as explanatory variables. Trade costs include tariffs and shipping costs; and shipping costs include freight, insurance, and other charges. Since tariff is invariant in this case, the trade costs include only shipping costs. Moreover, the model does not include distance variable, which is commonly used in measuring shipping costs, because Japan and Korea are located similar distance from the United States. The main hypothesis tested in this paper is that as shipping costs increase, U.S. imports of higher-quality cars from Japan and Korea increase.

[12] Import shares are calculated with data from the United States International Trade Commission (USITC, 2009).

The model proposed in this paper utilizes freight charge, the main part of shipping costs, as an increasing function of the price of oil based on an observation that the price of oil positively affects shipping rates with a one-month lag (Kilian, 2009). Shipping costs include other charges such as insurance and handling. The inference is that freight charge measures per-unit shipping cost, whereas insurance and other charges may vary with car prices. Thus, higher insurance and other charges reduce or eliminate the Alchian-Allen effect on the demand for higher quality cars.

The nominal exchange rate is utilized in the present model. Given currency supply and demand conditions, an appreciation of the dollar is expected to lower the price of foreign goods and increase the level of exports. An indicator variable for the Asian financial crisis of 1997-1998 is used to control for the large depreciation of the Korean won against the dollar.

This paper argues that the Alchian-Allen theorem holds when shipping costs account for per-unit freight charge only without considering insurance and other charges that may be imposed based on the car price. Another condition for the Alchian-Allen theorem to hold is that the effect of per-unit charge is larger than the effect of the ad-valorem charges. Present estimates based on Baltic Dry shipping index (and the price of crude oil) as a measure of freight rate support the Alchian-Allen theorem. Estimations based on actual shipping cost data are also conducted, and the result indicates that the effect of insurance and other charges outweighs the effect of per-unit freight charge.[13] It suggests that actual shipping cost data do not support the Alchian-Allen theorem. The relative price of higher quality cars is not necessarily lower when shipping costs rise, perhaps due to higher insurance and other charges. Furthermore, this paper

[13] Actual shipping cost is calculated by deducting the Custom value from the CIF (cost, insurance, and freight) value, and the result is divided by the total quantity of car imports. The Custom value is defined by the United States International Trade Commission (USITC, 2009) as “the price actually paid or payable for merchandise, excluding U.S. import duties, freight, insurance, and other charges.” The CIF value excludes U.S. import duties.

suggests that foreign car production is to be done domestically when trade costs are high, regardless differences in car quality.

The remainder of the paper proceeds as follows. Section 2 provides literature review. Section 3 discusses the data. Section 4 describes the shipping model, and Section 5 discusses the empirical method and results. Section 6 concludes the paper.

3.2 Literature review

Alchian and Allen (1964) propose that a per-unit charge applied to two similar goods will lower the relative price and increase the relative consumption of the higher quality good (e.g., Hummels and Skiba [2002] and Bauman [2004]). This theorem, often called "shipping the good apples out," has been proven algebraically by Borcherding and Silberberg (1978) and Saito (2007). The theorem originally applies to a two-good world only, but Borcherding and Silberberg (1978) subsequently provide a proof of its application with two goods in a many-good world given the two goods are close substitutes. In a more recent paper, Bauman (2004) relaxes the assumption of close substitutability between two goods in a many-good world.

A recent application of "shipping the good apples out" in international trade is reported by Hummels and Skiba (2002, 2004), who provide theoretical and empirical evidence based on bilateral trade data for six countries and their trading partners. With data on traded good prices, quantities, and shipping costs, they point out that shipping costs behave like a quantitative restriction similar to quotas (2002, p. 4). Another important contribution of Hummels and Skiba is their analysis of variation in ad-valorem trade costs such as tariffs, suggesting that they lower the relative demand for high quality goods and reduce the Alchian-Allen effect.

Issues on automobile trade between Japan and the United States have been on the political agenda of the two countries for a few decades. In 1981, Japan voluntarily limited automobile exports to the United States with a three-year export quota of 1.68 million automobiles (Tharp, 1981). The effectiveness of the voluntary export restraint (VER) system in initiating recovery of the U.S. automobile industry was in question as Japan shifted its concentration to exporting larger and higher-priced cars, increasing its export revenue while complying with the quantity restriction. The result was increased competition in the U.S. higher-priced automobile market dominated by U.S. automakers. A trade dispute occurred from 1993 to 1995. An agreement ended the two-year dispute after the Japanese agreed to deregulate its domestic automobile market. Another important deal reached in 1995 was that Japanese automobile manufacturers in the United States would increase production by 25% to 2.65 million automobiles in 1998 (Katzner and Nikomarov, 2008).

Kilian (2009) notes that freight rates may increase during upward oil price shocks due to higher demand for commodities and the use of bunker fuel oil as an input in shipping services. Kilian finds that the changes in freight rates do not occur contemporaneously with the changes in the price of crude oil and suggests a one-month lag on the response of freight rates. Backus and Crucini (2000) report that there is a relationship between the prices of oil and the terms of trade similar to the relationship between nominal or real exchange rates and the terms of trade. Their finding suggests that the price of oil and exchange rate are important determinants of trade.

Banik and Biswas (2007) point out that price competition among firms determines the degree of exchange rate pass-through in the U.S. automobile market with the degree of price competition negatively correlated with the degree of exchange rate pass-through. They find low degree of exchange rate pass-through for the Japanese and Korean automobile exporters, indicating that exporters from these two countries try to protect their market share in the U.S. market by offsetting the effect of exchange rate changes on car prices.

Korean won experienced a large depreciation against the dollar during the Asian financial crisis of 1997-1998. Korea was one of the Asian countries hardest hit by the crisis, along with Indonesia and Thailand. A rapid increase of short-term foreign borrowings before the crisis, combined with incomplete financial sector reforms, are identified as the main problems leading to corporate bankruptcies and financial sector collapses in these countries when the exchange rate movements turned against them (Radelet and Sachs, 1998). The Asian financial crisis did not significantly affect the Japanese economy, and the yen mildly depreciated against the dollar.

3.3 Data

A time-series model is applied to monthly data of U.S. car imports from Japan and Korea to test the main hypothesis on consistency with the Alchian-Allen theorem. Four data sets are constructed for the period 1995 to 2005 with 132 observations in each. The four data sets represent higher and lower quality cars from the two exporting countries. Two types of automobiles are examined: passenger motor vehicles with engine sizes exceeding 1,000 cc but not exceeding 1,500 cc (small size cars), and passenger motor vehicles with engine sizes exceeding 1,500 cc but not exceeding 3,000 cc (medium size cars). In addition, two data sets of higher quality cars from 1995 and 2008 with 180 observations are examined to include more recent data. [14]

[14] The lower quality car imports (small size cars) from Korea discontinued after 2005.

The data on car imports (in quantities and values) are obtained from the Interactive Tariff and Trade Database of the United States International Trade Commission (USITC, 2009). The dollar exchange rates are based on noon buying rates for JPY/USD and KRW/USD reported by the Federal Reserve Bank of New York (2009). The shipping index used is the Baltic Dry Index, which is taken from the FactSet Research Systems database. The data on crude oil price is the WTI spot price FOB taken from the Energy Information Administration of the U.S. Department of Energy (2009). All prices are deflated using Consumer Price Index reported by the Bureau of Labor Statistics (2009).

U.S. imports of small size cars from Japan exhibits an increasing trend during the last five years (Figure 3.1). In the same period, small size cars from Korea show a declining trend. In medium size category, car imports from Japan decreased during the period 1995-2003 and then increased during the last five years of observations, whereas car imports from Korea increased steadily over the period 1995-2008. The unit price of different car categories are presented in Figure 3.2, showing a persistent increase in prices for both small cars from Japan and Korea during the period 2000 to 2005. Further, Figure 3.3 shows actual shipping costs (CIF charges) for both small and medium size cars.

Other data series are presented in Figure 3.4. The price of crude oil exhibits an increasing trend prior to 2006. After a short decline in 2006, it shows a rapid increase in 2007-2008, with a peak in the second quarter of 2008. The Baltic Dry Index was relatively stable until 2003, and then considerably higher and more volatile through 2008. Issued daily by the London-based Baltic Exchange, the index measures worldwide international shipping rates of many dry bulk cargoes. The plots of crude oil price and Baltic Dry Index indicate a delay on the response of shipping rates to changes in oil price. The exchange rate data indicate that the Korean won experienced a large depreciation against the dollar during the Asian financial crisis of 1997-1998. The Japanese yen value was relatively stable over the period.

[Figures 3.1 through 3.4 here]

3.4 The shipping model

The model estimates the quantity of U.S. car imports from Japan and Korea with three main explanatory variables: car price, shipping costs, and the exchange rate. All variables in the model are in natural log form.

$$c_{xi} = F(p_{xi}, z_{xi}, e_i) \tag{1}$$

where subscript x denotes small car (s) or medium car (m), and subscript i denotes exporting country, Japan (j) or Korea (k). The variables are defined as follows:

c = quantity of car imports

p = car price per unit based on the U.S. Custom value (excluding shipping costs)

z = shipping costs, measured in two ways:

(a) using freight rate (f), constructed with Baltic Dry Index as an increasing function of oil price (o), where $f = F(o_{t-1})$.

(b) using actual shipping costs(t_{xi}), calculated from the U.S. Custom value and CIF value of imported car.

e = dollar exchange rate (USD/JPY or USD/KRW) with an indicator variable controlling for the Asian financial crisis in the case of KRW.

[Table 3.1 here]

The relative quality of each import category is determined by comparing its average price relative to those of other import categories as shown in Table 3.1. Average car prices are calculated based on the U.S. Custom value, which is the price actually paid or payable for merchandise, excluding U.S. import duties, freight, insurance, and other charges (USITC, 2010). The average prices of the four import categories can be compared as follows. The cars c_{sj} have a higher price than c_{sk}, but the cars c_{mj} has a lower price than c_{mk}.

The present study assumes similar product variations in c_{sj} and c_{sk}, and applies the same assumption to c_{mj} and c_{mk}. Based on the comparison of car prices in Table 3.1, the relative quality of each import category can be determined: *higher quality goods* are small size cars from Japan and medium size cars from Korea, and *lower quality goods* are medium size cars from Japan and small size cars from Korea. To be clear, higher quality cars are marked as c_{sj}^h and c_{mk}^h, and lower quality cars are marked as c_{sk}^l and c_{mj}^l. A detailed illustration of the Alchian-Allen theorem is to be explained in the following paragraphs.

A world with n goods comprises c_{sj}^h, c_{sk}^l, c_{mj}^l, c_{mk}^h, and y, where y is a composite of "other" goods. The first and second goods are small cars with different qualities, higher and lower. The third and fourth goods are medium cars with higher quality and lower quality. Higher quality cars are represented by c_h and lower quality cars are represented by c_l; then by assumption, the car prices follow $p_h > p_l > 0$.

A per-unit shipping cost (freight charge), f, is added to the prices of c_h and c_l. The prices become $p_h + f$, $p_l + f$, and p_y, respectively.

Following Borcherding and Silberberg (1978),

$$\partial(c_h/c_l)/\partial f > 0 \tag{2}$$

where $c_h(p_h, p_l, p_y, U)$ and $c_l(p_h, p_l, p_y, U)$ are Hicksian demand functions. That is, as the shipping cost rises, the consumption of higher quality cars increases relative to the lower quality cars, holding real income and the prices of all other goods constant.

Given $\varepsilon_{ij} = (p_j/c_i)\ (\partial c_i/\partial p_j)$ be the compensated elasticities, applying the chain rule and quotient rule on (2) will result in

$$\partial(c_h/c_l)/\partial f \ = \ \frac{c_h}{c_l}\left[(\varepsilon_{hh}/p_h) \ + \ (\varepsilon_{hl}/p_l) - (\varepsilon_{lh}/p_h) - (\varepsilon_{ll}/p_l)\right] > \ 0 \qquad (3)$$

which confirms the Alchian-Allen theorem.

In a two-good world, Borcherding and Silberberg (1978) note that Hick's third law (1946) applies ($\sum_i \ \varepsilon_{ij} = \ 0$). This implies that the two cars in a two-car world must be substitutes ($\varepsilon_{hl} > 0$) and own-price elasticities are negative ($\varepsilon_{ll} < 0$), which can be obtained by substituting for ε_{hh} and ε_{lh} (given $\varepsilon_{hl} = -\varepsilon_{lh}$). In an *n*-good world (introducing a third, composite of "other" goods), Borcherding and Silberberg show that Alchian-Allen holds if the two goods are assumed to be close substitutes.

Bauman (2004) relaxes the assumption of close substitutability between the two goods in an *n*-good world. So long as the two goods are not close complements ($\varepsilon_{hl} > \varepsilon_{ll}$) and are not close in price ($p_h >> p_l$) a per-unit charge will increase the relative consumption of the higher quality good (2004, p.535). The implication is that the two goods bearing per-unit charge do not have to be close substitutes. In the car import case, higher quality cars from Japan (c_{sj}^h) can be substitute to lower quality cars from Japan (c_{mj}^l) or lower quality cars from Korea (c_{sk}^l). When per-unit shipping costs increase, U.S. demand for higher quality cars from Japan increases relative to lower quality cars from both Japan and Korea.

The consumer utility is based on Hummels and Skiba (2002) and applied specifically to the utility obtained by the U.S. consumer from buying imported cars. The U.S. consumer obtains greater utility from higher quality cars,

$$U_x = \left[\sum_{x,i} \lambda_{xi}^{l}\left(c_{xi}^{l}\right)^{\theta} + \lambda_{xi}^{h}\left(c_{xi}^{h}\right)^{\theta}\right]^{1/\theta}, \tag{4}$$

denoting c as car quantity and λ as car quality, where $\lambda_{xi}^{h} > \lambda_{xi}^{l}$ and $\theta = 1 - \frac{1}{\sigma}$, and σ is the elasticity of substitution between higher and lower quality cars based on the CES utility function. Subscript x denotes small car (s) or medium car (m), and subscript i denotes exporting country, Japan (j) or Korea (k).

The price of cars depends on the U.S. customer price (p_{xi}) and trade costs. Trade costs include tariffs and shipping costs; and shipping costs include freight, insurance, and other charges. Since tariff is invariant in this case, the trade costs include only shipping costs (z).

Let f denotes freight rate and g denotes ad-valorem insurance and other shipping charges, then the price of car at destination follows

$$p_{xi}^{US} = f + g_{xi}\, p_{xi} \tag{5}$$

where $g > 1$.

Freight rate f represents a per-unit cost that follows the Alchian-Allen theorem, assumed the same for Japan and Korea given similar distance to the United States; and g is ad-valorem charges that increase with car price, $g = F(p)$. As described in (1), f increases with oil price (o_{t-1}). The model in (1) does not count for ad-valorem charges (g) when freight rate f is used, based on Baltic Dry Index and the price of oil. The ad-valorem charges g are taken into account when the model utilizes the actual CIF (cost, insurance, and freight) shipping costs instead of freight rate. The actual data are calculated based on the U.S. Custom value and CIF value of car imports as well as import quantity reported by the USITC (2010). The actual data do not allow for decomposing the CIF charges into cost, insurance, and freight separately.

Hummels and Skiba (2002) note that ad-valorem charges reduce the Alchian-Allen effect on shipping cost. Consequently, Alchian-Allen theorem is expected to hold when shipping costs do not count for ad-valorem charges, or when per-unit charge is larger than the ad-valorem charges. The present model assumes that the effect of insurance and other charges is similar to that of ad-valorem tariffs in reducing the Alchian-Allen effect.

An expenditure minimization problem for U.S. consumer buying imported cars can be constructed with the car price in (5) and consumer utility in (4). Assuming competitive firms, the consumer's first order conditions implies

$$\frac{c_{xi}^h}{c_{xi}^l} = \left(\frac{p_{xi}^l g_{xi} + f}{p_{xi}^h g_{xi} + f}\right)^\sigma \left(\frac{\lambda_{xi}^h}{\lambda_{xi}^l}\right)^\sigma. \tag{6}$$

Rearranging (6) results in

$$\frac{c_{xi}^h}{c_{xi}^l} = \left(\frac{p_{xi}^l + f/g_{xi}}{p_{xi}^h + f/g_{xi}}\right)^\sigma \left(\frac{\lambda_{xi}^h}{\lambda_{xi}^l}\right)^\sigma. \tag{7}$$

Further derivations of (7) with respect to f and g will give results that suggest the relative consumption of higher quality cars increases with shipping costs when the effect of per-unit freight charge is larger than the effect of ad-valorem insurance and other charges.

3.5 Empirical method and results

The augmented Dickey-Fuller (ADF) test for the stationarity of individual variables is summarized in Table 3.2. The actual shipping cost variables are all stationary in their log forms. Other variables are mostly difference-stationary in their log forms. The spurious regressions are presented in Table 3.3 and Table 3.4. Two regressions are run for each data set: the first includes freight rate (f), whereas the second includes actual shipping costs (t) instead. Estimations using freight rate are cointegrated by the Engle-Granger (EG) for all four data sets, c_{sj}^{h}, c_{sk}^{l}, c_{mj}^{l}, and c_{mk}^{h} (Table 3.3[a]). Estimations using actual shipping costs are cointegrated for three data sets (c_{sj}^{h}, c_{sk}^{l}, and c_{mk}^{h}), but not cointegrated for c_{mj}^{l} (Table 3.3[b]). The ARCH test results indicate that heteroskedasticity exists in the data, and the Breusch-Godfrey Lagrange Multiplier test results indicate the presence of autocorrelations.

[Tables 3.2 through 3.4 here]

The model is estimated using the dynamic OLS (DOLS) regression proposed by Stock and Watson (1993). It estimates long-run equilibria in cointegrated systems that may have variables integrated with different orders. The DOLS estimates parameter β of the following regression:

$$y_{1t} = \beta' y_{2t} + d(L)\mathit{\Delta} y_{2t} + v_t \qquad (8)$$

where $d(L)$ is the leads and lags of $\mathit{\Delta} y_{2t}$, the first differences of any $I(1)$ variables; and v_t is a stochastic error term. This process eliminates asymptotically possible bias caused by endogeneity or serial correlation.

DOLS is applied to the shipping model as follows,

$$c_{xit} = \beta_1' p_{xit} + d(L)\mathit{\Delta} p_{xit} + \beta_2' z_{xit} + d(L)\mathit{\Delta} z_{xit} + \beta_3' e_{it} + d(L)\mathit{\Delta} e_{it} + v_t \qquad (9)$$

substituting $$z_{xit} = f_t \tag{9a}$$

or $$z_{xit} = t_{xit} \tag{9b}$$

where freight rate (f_t) and actual shipping costs (t_{xit}) are the variables of interest. The estimation results are discussed separately based on (9a) and (9b).

[Tables 3.5 and 3.6 here]

(a) Regressions with freight rate

Estimations based on monthly data from 1995 to 2005, using freight rate (Baltic Dry Index and the price of oil) as a measure of shipping cost, show expected results on the response of import quantity to the changes in freight rate (Table 3.5[a]). An increase in freight rate is related to an increase in U.S. imports of higher quality cars from Japan and Korea (c_{sj}^h and c_{mk}^h) and a decrease in U.S. imports of lower quality cars from Japan (c_{mj}^l), holding car price and exchange rate constant. These results are significant at the 1% level of significance. The coefficient estimate for lower quality cars from Korea (c_{sk}^l) is not significant.

Exchange rate variable shows the expected significant negative coefficients for three data sets (c_{sj}^h, c_{sk}^l,and c_{mk}^h) and negative but not significant coefficient for c_{mj}^l. A depreciation of the U.S. dollar increases the foreign car price and thus lowers the demand for those cars, reflected in the lower quantity of car imports.

Extending the period of observations to 2008 (Table 3.6[a]), the regression results for higher quality cars show similar results on the coefficient estimates of the freight rate variable, but weaker in magnitudes. Still, it seems that the model works even in the period of oil price shock in 2006 to 2008.

(b) Regressions with actual shipping costs

Estimations based on monthly data from 1995 to 2005 using actual CIF shipping charges as a measure of shipping costs are presented in Table 3.5[b]. The significant negative coefficient of shipping cost variable for the U.S. import of higher quality cars from Korea (c_{mk}^{h}) implies that an increase in shipping costs is related to a decrease in import quantity of higher quality cars, holding car price and exchange rate constant. This result seems to counter the Alchian-Allen theorem. Similarly, the significant positive coefficient of shipping cost variable for the lower quality cars from Korea (c_{sk}^{l}) shows inconsistency with Alchian-Allen theorem.

The exchange rate variable shows significant negative coefficients consistent with the results from (a) for two data sets (c_{sk}^{l} and c_{mj}^{l}). The exchange rate variable for higher quality cars from Japan has a significant positive coefficient, suggesting that a lower U.S. dollar value, relative to the Japanese yen, increases the import quantity of higher quality cars from Japan. This result counters the result from (a).

Extending the period of observations to 2008 (Table 3.6[b]), the regression results for higher quality cars from Japan and Korea (c_{sj}^{h} and c_{mk}^{h}) show coefficient estimates similar to those from the data sets of shorter period of observations (1995 to 2005).

3.6 Conclusion

This paper examines data on Japanese and Korean automobile exports to the United States to determine consistency with the Alchian-Allen theorem using the dynamic OLS estimators. The result suggests that the Alchian-Allen theorem holds when shipping costs account for per-unit freight charge only without considering insurance and other shipping charges that may be imposed based on the car price. It also suggests that another condition for Alchian-Allen theorem to hold is when the effect of per-unit charge is larger than the effect of the ad-valorem charges. Hummels and Skiba (2002) report that ad-valorem trade costs such as tarrifs reduce the Alchian-Allen effect. Similarly, the present study finds that insurance and other shipping charges seem to reduce or eliminate the Alchian-Allen effect.

Estimations based on Baltic Dry shipping index and the price of oil as a measure of freight rate support the Alchian-Allen theorem. As shipping rates increase, U.S. imports of higher quality cars from Japan and Korea increase relative to the lower quality cars. In contrast, estimations based on the actual CIF (cost, insurance, and freight) charges indicate that the relative price of higher-quality car imports is not necessarily lower when shipping costs rise. A possible rationale is that insurance and other shipping charges are imposed based on the car price. These charges reduce or eliminate the Alchian-Allen effect of the per-unit freight charge. Therefore, the present shipping cost data do not support the Alchian-Allen theorem.

The model seems to work better for car imports from Korea than for car imports from Japan, as indicated by the cointegration test and DOLS regression results. A possible explanation is there are more product variations in the cars imported from Japan compared to those imported from Korea. Another possible explanation is there are more automobile trade agreements between Japan and the United States that may influence the types of cars imported from Japan.

The result of this study suggests that foreign car production is to be done domestically in the United States when trade costs are high, regardless differences in car quality. Policy that promotes domestic production by foreign car manufacturers is recommended.

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APPENDIX 1

Tables and Figures for Chapter 1

Figure 1.1. Sources of funds for a corporation

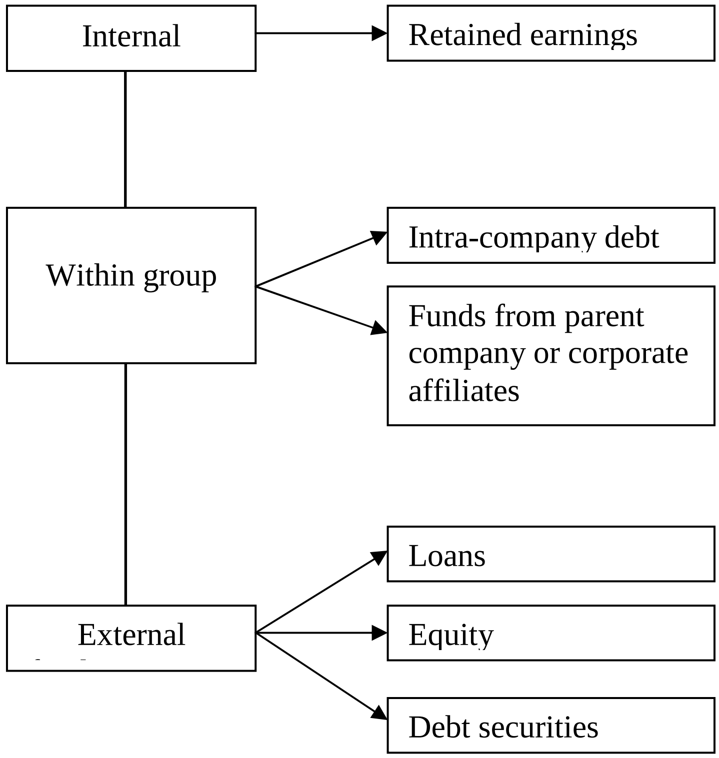


Table 1.1. List of countries

Main countries	Partner countries
Germany	Algeria
Japan	Australia
United States	Austria
	Belgium
	Bulgaria
	Canada
	Chile
	China
	Colombia
	Costa Rica
	Croatia
	Czech Republic
	Denmark
	Dominican Republic
	Ecuador
	Finland
	France
	Germany
	Greece
	Hungary
	Ireland
	Israel
	Italy
	Japan
	Luxembourg
	Malaysia
	Morocco
	Netherlands
	New Zealand
	Nigeria
	Norway
	Pakistan
	Philippines
	Poland
	Portugal
	Romania
	Russian Federation
	Singapore
	Slovak Republic
	South Africa
	Spain
	Sweden
	Switzerland
	Tunisia
	Ukraine
	United Kingdom
	United States
	Venezuela

Table 1.2. Variable descriptions and summary statistics

(a) Variable descriptions

Variable	Variable label	Description
TRA_{ijt}	ln[(xij+xji)/(gdpi*gdpj)]	xij+xji is trade between country i and country j in million US$ gdpi(j) is GDP of country i(j) in billion US$
DIS_{ij}	ln(disij)	disij is distance between country i and country j in kilometers
LCB_{ij}	land common border	1 = share common border, 0 = no common border
DEV_{ijt}	stage of development	1 = same stage, 0 = different stage
LNA_{ijt}	ln(lnai*lnaj)	lnai(j) is score of access to loan of country i(j)
EQA_{ijt}	ln(eqai*eqaj)	eqai(j) is score of access to equity of country i(j)
CCR_{ijt}	ln(max ccri,ccrj/min ccri,ccrj)	ccri(j) is country credit rating of country i(j)
CAP_{ijt}	ln(capi*capj)	capi(j) is score of capital controls in country i(j)
RER_{ijt}	ln(max reri,rerj/min reri,rerj)	reri(j) is real effective exchange rate in country i(j)

(b) Summary statistics

Variable	Observation	Mean	Standard Deviation	Minimum	Maximum
TRA_{ijt}	690	-5.036	1.336	-7.621	-1.754
DIS_{ij}	690	8.540	1.020	5.156	9.843
LCB_{ij}	690	0.0725	0.259	0	1
DEV_{ijt}	690	0.4609	0.499	0	1
LNA_{ijt}	690	2.681	0.339	1.658	3.352
EQA_{ijt}	690	3.395	0.229	2.591	3.728
CCR_{ijt}	690	0.346	0.350	0	1.525
CAP_{ijt}	690	3.642	0.375	2.480	4.345
RER_{ijt}	690	0.210	0.170	0.000	0.953

Table 1.3. Pair-wise correlations between variables

	TRA_{ijt}	DIS_{ij}	LCB_{ij}	DEV_{ijt}	LNA_{ijt}	EQA_{ijt}	CCR_{ijt}	CAP_{ijt}	RER_{ijt}
TRA_{ijt}	1								
DIS_{ij}	-0.702*	1							
	(0.000)								
LCB_{ij}	0.479*	-0.679*	1						
	(0.000)	(0.000)							
DEV_{ijt}	0.034	-0.156*	0.190*	1					
	(0.376)	(0.000)	(0.000)						
LNA_{ijt}	0.089*	-0.139*	0.122*	0.551*	1				
	(0.019)	(0.000)	(0.001)	(0.000)					
EQA_{ijt}	0.066	0.039	0.087*	0.567*	0.489*	1			
	(0.082)	(0.301)	(0.022)	(0.000)	(0.000)				
CCR_{ijt}	-0.064	0.171*	-0.206*	-0.725*	-0.578*	-0.568*	1		
	(0.092)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)			
CAP_{ijt}	0.117*	-0.118*	0.134*	0.504*	0.240*	0.278*	-0.358*	1	
	(0.002)	(0.002)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		
RER_{ijt}	-0.359*	0.361*	-0.244*	-0.061	-0.115*	-0.118*	0.033	-0.101*	1
	(0.000)	(0.000)	(0.000)	(0.110)	(0.003)	(0.002)	(0.387)	(0.008)	

* indicates correlation is significant at 5% level. P-values are in parentheses.

Table 1.4. Regression results

(a) Country-specific and country-pair time fixed effects[†]

Main countries: United States, Japan, and Germany.	TRA_{ijt}	Partner countries (1) All partner countries	(2) Developed countries	(3) Less developed countries
Gravity	DIS_{ij}	-0.643***	-0.483***	-0.794***
		(0.038)	(0.053)	(0.053)
	LCB_{ij}	0.023	0.570***	0.116
		(0.113)	(0.133)	(0.201)
	DEV_{ijt}	-1.321***	---	---
		(0.120)		
External Funds	LNA_{ijt}	0.735***	0.808**	1.097***
		(0.205)	(0.315)	(0.258)
	EQA_{ijt}	0.425**	-1.839***	0.453**
		(0.173)	(0.504)	(0.196)
International Finance	CCR_{ijt}	-0.883***	-1.800***	-0.622***
		(0.157)	(0.531)	(0.181)
	CAP_{ijt}	0.548***	-1.209***	0.588***
		(0.111)	(0.382)	(0.123)
	RER_{ijt}	-0.292	2.779***	-0.776**
		(0.286)	(0.686)	(0.327)
Intercept		-3.349***	8.927***	-3.267***
		(0.665)	(2.320)	(0.782)
Observations		690	318	372
Adjusted R^2		0.937	0.958	0.936

[†] Anderson and van Wincoop (2003) and Baier and Bergstrand (2007), selected model.
***,**,* indicate variable is significant at 1%, 5%, and 10% levels, respectively. Standard errors are in parentheses.
--- indicates variable is omitted.

(b) Country-specific time fixed effects††

Main countries: United States, Japan, and Germany.	TRA_{ijt}	Partner countries (1) All partner countries	(2) Developed countries	(3) Less developed countries
Gravity	DIS_{ij}	-1.100***	-0.903***	-1.314***
		(0.051)	(0.070)	(0.068)
	LCB_{ij}	-0.335*	-0.092	0.467
		(0.172)	(0.197)	(0.313)
	DEV_{ijt}	-0.525**	---	---
		(0.214)		
External Funds	LNA_{ijt}	0.216	0.02	0.395
		(0.222)	(0.297)	(0.300)
	EQA_{ijt}	1.129***	2.852***	0.924***
		(0.294)	(0.000)	(0.328)
International Finance	CCR_{ijt}	0.442*	0.912	0.311
		(0.245)	(0.721)	(0.268)
	CAP_{ijt}	0.264	0.240	0.385*
		(0.205)	(0.564)	(0.229)
	RER_{ijt}	-0.465	-0.279	-0.867*
		(0.365)	(0.585)	(0.458)
Intercept		-0.683	-8.313**	1.098
		(1.110)	(3.765)	(1.246)
Observations		690	318	372
Adjusted R^2		0.700	0.781	0.698

††Anderson and van Wincoop (2003), ignoring country-pair fixed effects.
***,**,* indicate variable is significant at 1%, 5%, and 10% levels, respectively. Standard errors are in parentheses.
--- indicates variable is omitted.

(c) OLS†††

Main countries: United States, Japan, and Germany.	TRA_{ijt}	Partner countries (1) All partner countries	(2) Developed countries	(3) Less developed countries
Gravity	DIS_{ij}	-0.937***	-0.763***	-1.129***
		(0.050)	(0.068)	(0.072)
	LCB_{ij}	-0.030	0.263	0.307
		(0.185)	(0.215)	(0.348)
	DEV_{ijt}	-0.531***	---	---
		(0.115)		
External Funds	LNA_{ijt}	0.029	-0.046	0.078
		(0.132)	(0.183)	(0.186)
	EQA_{ijt}	1.222***	1.693***	1.187***
		(0.203)	(0.448)	(0.239)
International Finance	CCR_{ijt}	0.272*	0.220	0.332*
		(0.157)	(0.485)	(0.179)
	CAP_{ijt}	0.321***	0.302	0.357***
		(0.107)	(0.191)	(0.136)
	RER_{ijt}	-0.648***	-0.614	-0.856***
		(0.224)	(0.374)	(0.293)
Intercept		-2.147**	-5.545***	-0.608
		(0.862)	(1.750)	(1.075)
Observations		690	318	372
Adjusted R^2		0.537	0.621	0.499

†††Ordinary Least Squares, ignoring fixed effects.
***,**,* indicate variable is significant at 1%, 5%, and 10% levels, respectively. Standard errors are in parentheses.
--- indicates variable is omitted.

Table 1.5. Fixed effects regressions without pair-country interactions for the financial development variables

Main countries: United States, Japan, and Germany.	TRA_{ijt}	Partner countries (1) All partner countries	(2) Developed countries	(3) Less developed countries
Gravity	DIS_{ij}	-0.548***	-0.419***	-0.585***
		(0.034)	(0.061)	(0.046)
	LCB_{ij}	0.095	0.519***	0.144
		(0.099)	(0.140)	(0.147)
	DEV_{ijt}	-1.253***	---	---
		(0.108)		
External Funds	LNA_{it}	0.990	6.867***	1.056
		(0.999)	(0.692)	(0.916)
	LNA_{jt}	0.074	0.849**	0.423
		(0.236)	(0.352)	(0.326)
	EQA_{it}	-3.402	---	-3.068
		(2.156)		(1.977)
	EQA_{jt}	0.320*	-1.871***	0.455***
		(0.169)	(0.591)	(0.167)
International Finance	CCR_{it}	6.921*	---	7.466**
		(3.607)		(3.340)
	CCR_{jt}	1.280***	0.642	1.259***
		(0.149)	(0.522)	(0.165)
	CAP_{it}	3.641***	---	3.602***
		(0.632)		(0.595)
	CAP_{jt}	0.154	-0.990**	0.099
		(0.120)	(0.391)	(0.125)
	RER_{it}	-2.291	---	-2.730*
		(1.598)		(1.506)
	RER_{jt}	1.761***	-0.772	2.268***
		(0.222)	(0.619)	(0.249)
Intercept		-36.46***	-4.69	-39.824***
		(11.75)	(3.114)	(10.827)
Observations		690	318	372
Adjusted R^2		0.955	0.954	0.967

***,**,* indicate variable is significant at 1%, 5%, and 10% levels, respectively. Standard errors are in parentheses.
--- indicates variable is omitted.

APPENDIX 2

Tables and Figures for Chapter 2

Table 2.1. Commodity price data description

Category	ID	Commodity	IMF Code
Metals	1	Aluminum, LME standard grade, minimum purity, CIF UK	PALUM
	2	Copper, LME, grade A cathodes, CIF Europe	PCOPP
	3	Iron Ore Carajas	PIORECR
	4	Lead, LME, 99.97 percent pure, CIF European	PLEAD
	5	Nickel, LME, melting grade, CIF N Europe	PNICK
	6	Tin, LME, standard grade, CIF European	PTIN
	7	Zinc, LME, high grade, CIF UK	PZINC
	8	Uranium, NUEXCO, Restricted Price, US$ per pound	PURAN
Fuels	9	Coal thermal for export, Australia	PCOALAU
	10	Oil, Average of U.K. Brent, Dubai, and West Texas Intermediate	POILAPSP
	11	Oil, UK Brent, light blend 38 API, fob U.K.	POILBRE
	12	Oil, Dubai, medium, Fateh 32 API, fob Dubai	POILDUB
	13	Oil, West Texas Intermediate, 40 API, Midland Texas	POILWTI
	14	Natural Gas, BEA	
Food	15	Bananas, avg of Chiquita, Del Monte, Dole, U.S. Gulf delivery	PBANSOP
	16	Barley, Canadian Western No. 1 Spot	PBARL
	17	Beef, Australia/New Zealand frozen, U.S. import price	PBEEF
	18	Cocoa, ICO price, CIF U.S. & European ports	PCOCO
	19	Coconut Oil, Philippines/Indonesia, CIF Rotterdam	PROIL
	20	Fishmeal, 64/65 percent, any orig, CIF Rotterdam	PFISH
	21	Groundnut, U.S. runners, CIF European	PGNUTS
	22	Lamb, New Zealand, PL frozen, London price	PLAMB
	23	Maize, U.S. number 2 yellow, fob Gulf of Mexico	PMAIZMT
	24	Olive Oil, less that 1.5% FFA	POLVOIL
	25	Orange Brazilian, CIF France	PORANG
	26	Palm Oil, Malaysia and Indonesian, CIF NW Europe	PPOIL
	27	Hogs, 51-52% lean, 170-191 lbs, IL, IN, OH, MI, KY	PPORK
	28	Chicken, Ready-to-cook, whole, iced, FOB Georgia Docks	PPOULT
	29	Rice, 5 percent broken, nominal price quote, fob Bangkok	PRICENPQ
	30	Norwegian Fresh Salmon, farm bred, export price	PSALM
	31	Shrimp, U.S., frozen 26/30 count, wholesale NY	PSHRI
	32	Soybean Meal, 44 percent, CIF Rotterdam	PSMEA
	33	Soybean Oil, Dutch, fob ex-mill	PSOIL
	34	Soybean, U.S., CIF Rotterdam	PSOYB
	35	Sugar, EC import price, CIF European	PSUGAEEC
	36	Sugar, International Sugar Agreement price	PSUGAISA
	37	Sugar, U.S., import price contract number 14 CIF	PSUGAUSA
	38	Sunflower Oil, any origin, ex-tank Rotterdam	PSUNO
	39	Wheat, U.S. number 1 HRW, fob Gulf of Mexico	PWHEAMT
Beverages	40	Coffee, Other Milds, El Salvdor and Guatemala, ex-dock New York	PCOFFOTM
	41	Coffee, Robusta, Uganda and Cote dIvoire, ex-dock New York	PCOFFROB
	42	Tea, From July 1998,Kenya auctions, Best Pekoe Fannings. Prior, London auctions, CIF U.K. warehouses	PTEA
Raw Materials	43	Cotton, Liverpool Index A, CIF Liverpool	PCOTTIND
	44	Wool Coarse, 23 micron, AWEX	PWOOLC
	45	Wool Fine, 19 micron, AWEX	PWOOLF

Continue to next page.

Continued from previous page.

Category	ID	Commodity	IMF Code
Industrial Inputs	46	Hides, U.S., Chicago, fob Shipping Point	PHIDE
	47	Log, soft, export from U.S. Pacific coast	PLOGORE
	48	Log, hard, Sarawak, import price Japan	PLOGSK
	49	Rubber, Malaysian, fob Malaysia and Singapore	PRUBB
	50	Sawnwood, dark red meranti, select quality	PSAWMAL
	51	Sawnwood, average of softwoods, U.S. West coast	PSAWORE

Note: i) All data is obtained from IMF Primary Commodity Prices database with an exception of natural gas (ID#14). The U.S. wellhead natural gas data is obtained from the U.S. Energy Information Administration.

Table 2.2. Augmented Dickey-Fuller test and Cross-section Dependence test results: Nominal commodity prices

ID	*ADF*	*p*-value	ID	*ADF*	*p*-value	ID	*ADF*	*p*-value
1	-3.062*	0.026	18	-1.616	0.472	35	-1.641	0.456
2	-1.157	0.690	19	-1.826	0.359	36	-2.706	0.067
3	0.215	0.977	20	-1.212	0.666	37	-7.394*	0.000
4	-0.847	0.803	21	-4.047*	0.001	38	-2.699	0.069
5	-1.878	0.334	22	-2.207	0.197	39	-2.505	0.108
6	-2.078	0.246	23	-2.877*	0.043	40	-2.561	0.093
7	-2.104	0.229	24	-1.660	0.448	41	-2.061	0.246
8	-1.182	0.682	25	-1.998	0.278	42	-4.194*	0.000
9	-1.879	0.334	26	-2.996*	0.031	43	-3.487*	0.007
10	-0.912	0.787	27	-3.354*	0.011	44	-2.051	0.254
11	-0.999	0.754	28	-0.842	0.811	45	-3.412*	0.010
12	-0.818	0.819	29	-1.855	0.343	46	-3.181*	0.018
13	-1.048	0.738	30	-1.962	0.294	47	-1.550	0.504
14	-1.823	0.359	31	-2.608	0.085	48	-1.612	0.472
15	-2.940*	0.036	32	-3.043*	0.027	49	-1.367	0.601
16	-2.086	0.237	33	-2.654	0.077	50	-1.384	0.593
17	-1.961	0.294	34	-2.849*	0.048	51	-1.229	0.658

CD Statistic: 53.978, *p*-value: 0.000

Note: i) *ADF* denotes the augmented Dickey-Fuller t-statistic with an intercept. ii) Superscript * refers the case when the null of nonstationarity is rejected at the 5% significance level. iii) *CD* statistic is a cross-section dependence test statistic by Pesaran (2004) with the hypothesis of no cross-section dependence.

Table 2.3. Augmented Dickey-Fuller test and Cross-section Dependence test results: Relative commodity prices

ID	*ADF*	*p*-value	ID	*ADF*	*p*-value	ID	*ADF*	*p*-value
1	-3.569*	0.006	18	-2.689	0.071	35	-2.492	0.108
2	-2.087	0.237	19	-2.807	0.053	36	-3.274*	0.014
3	-0.969	0.763	20	-2.232	0.189	37	-2.946*	0.035
4	-1.843	0.351	21	-3.226*	0.016	38	-3.200*	0.017
5	-2.663	0.075	22	-3.805*	0.002	39	-2.706	0.067
6	-2.501	0.108	23	-2.636	0.080	40	-2.360	0.141
7	-2.740	0.063	24	-2.669	0.074	41	-2.035	0.262
8	-1.879	0.334	25	-3.751*	0.003	42	-3.036*	0.028
9	-2.410	0.133	26	-3.231*	0.016	43	-2.445	0.116
10	-1.864	0.343	27	-1.493	0.536	44	-2.590	0.088
11	-1.952	0.294	28	-5.692*	0.000	45	-2.364	0.141
12	-1.766	0.391	29	-2.540	0.097	46	-2.487	0.108
13	-2.080	0.246	30	-2.090	0.237	47	-1.735	0.407
14	-2.284	0.165	31	-0.709	0.843	48	-2.876*	0.043
15	-3.813*	0.002	32	-2.915*	0.040	49	-2.631	0.081
16	-3.870*	0.002	33	-2.705	0.068	50	-2.336	0.149
17	-2.146	0.213	34	-2.688	0.071	51	-2.237	0.181

CD Statistic: 48.313, *p*-value: 0.000

Note: i) *ADF* denotes the augmented Dickey-Fuller t-statistic with an intercept with the null of nonstationarity. ii) Superscript * refers the case when the null hypothesis is rejected at the 5% significance level. iii) *CD* statistic is a cross-section dependence test statistic by Pesaran (2004) with the null hypothesis of no cross-section dependence. iv) Each commodity price is deflated by the U.S. consumer price index to obtain the relative price.

Table 2.4. PANIC test results:
Nominal commodity prices

Idiosyncratic Components								
ID	*ADF*	*p*-value	ID	*ADF*	*p*-value	ID	*ADF*	*p*-value
1	-2.220*	0.024	18	-1.047	0.270	35	-1.083	0.254
2	-2.839*	0.004	19	-2.995*	0.003	36	-1.558	0.108
3	-0.952	0.302	20	-1.084	0.254	37	-4.079*	0.000
4	-1.324	0.173	21	-4.856*	0.000	38	-4.688*	0.000
5	-2.194*	0.026	22	-2.225*	0.024	39	-3.410*	0.001
6	-1.237	0.197	23	-2.372*	0.016	40	-2.258*	0.022
7	-2.298*	0.019	24	-1.438	0.141	41	-1.842	0.060
8	-0.845	0.351	25	-3.851*	0.000	42	-2.699*	0.006
9	-3.326*	0.001	26	-3.448*	0.001	43	-3.406*	0.001
10	-2.117*	0.030	27	-3.075*	0.002	44	-2.158*	0.027
11	-2.205*	0.025	28	-5.730*	0.000	45	-2.171*	0.027
12	-1.986*	0.042	29	-1.288	0.181	46	-1.414	0.149
13	-2.336*	0.018	30	-2.078*	0.033	47	-1.566	0.108
14	-1.577	0.108	31	-1.507	0.124	48	-2.240*	0.023
15	-5.327*	0.000	32	-2.719*	0.006	49	-1.448	0.141
16	-1.191	0.213	33	-1.708	0.082	50	-2.057*	0.035
17	-1.578	0.108	34	-1.860	0.057	51	-1.220	0.205

Panel Test Statistics: 21.445*, *p*-value: 0.000

Common Factor Components

ADF (Factor 1): -2.283, *p*-value: 0.165
ADF (Factor 2): -2.901*, *p*-value: 0.042

Note: i) *ADF* denotes the augmented Dickey-Fuller t-statistic with no deterministic terms (idiosyncratic components) and with an intercept (common factors) with the null hypothesis of nonstationarity. ii) Superscript * refers the case when the null hypothesis is rejected at the 5% significance level.

Table 2.5. PANIC test results:
Relative commodity prices

Idiosyncratic Components								
ID	*ADF*	*p*-value	ID	*ADF*	*p*-value	ID	*ADF*	*p*-value
1	-2.089*	0.032	18	-0.978	0.294	35	-1.311	0.173
2	-2.890*	0.004	19	-2.865*	0.004	36	-1.518	0.124
3	-0.798	0.367	20	-0.990	0.294	37	-3.429*	0.001
4	-1.396	0.157	21	-4.626*	0.000	38	-4.648*	0.000
5	-1.928*	0.048	22	-2.335*	0.018	39	-3.385*	0.001
6	-1.224	0.205	23	-2.233*	0.023	40	-2.078*	0.033
7	-2.201*	0.025	24	-1.658	0.091	41	-1.773	0.070
8	-0.807	0.367	25	-8.250*	0.000	42	-2.259*	0.022
9	-3.090*	0.002	26	-3.282*	0.001	43	-3.578*	0.001
10	-1.910*	0.050	27	-3.560*	0.001	44	-1.802	0.066
11	-1.974*	0.043	28	-2.269*	0.021	45	-2.316*	0.019
12	-2.062*	0.035	29	-1.114	0.246	46	-1.536	0.116
13	-2.197*	0.025	30	-2.038*	0.037	47	-1.666	0.090
14	-1.761	0.072	31	-1.868	0.056	48	-2.095*	0.031
15	-3.870*	0.000	32	-2.293*	0.020	49	-1.528	0.116
16	-1.071	0.262	33	-1.357	0.165	50	-2.214*	0.024
17	-1.170	0.221	34	-1.471	0.133	51	-1.671	0.089

Panel Test Statistics: 19.497*, *p*-value: 0.000

Common Factor Components

ADF (Factor 1): -1.887, *p*-value; 0.326
ADF (Factor 2): -2.912*, *p*-value: 0.040

Note: i) *ADF* denotes the augmented Dickey-Fuller t-statistic with no deterministic terms (idiosyncratic components) and with an intercept (common factors) with the null hypothesis of nonstationarity. ii) Superscript * refers the case when the null hypothesis is rejected at the 5% significance level. iii) Each commodity price is deflated by the U.S. consumer price index to obtain the relative price.

Table 2.6. Out-of-sample forecast performance:
Nominal commodity price factors

K	*RRMSPE*	*DMW*
1	1.0122	0.7039**
2	1.0026	0.2607
3	1.0054	0.5352*
4	1.0134	1.1514**

Note: i) Out-of-sample forecasting was recursively implemented by sequentially adding one additional observation from 180 initial observations toward 360 total observations. ii) *k* denotes the forecast horizon. iii) *RRMSPE* denotes the ratio of the root mean squared prediction error of the random walk hypothesis to the common factor model. iv) *DMW* denotes the test statistics of Diebold and Mariano (1995) and West (1996). v) * and ** denote rejection of the null hypothesis of equal predictability at the 10% and 5% significance levels, respectively. Critical values were obtained from McCracken (2007).

Table 2.7. Out-of-sample forecast performance:
Relative commodity price factors

K	*RRMSPE*	*DMW*
1	1.0169	0.9804**
2	1.0062	0.5834*
3	1.0082	0.7753**
4	1.0167	1.3815***

Note: i) Out-of-sample forecasting was recursively implemented by sequentially adding one additional observation from 180 initial observations toward 360 total observations. ii) *k* denotes the forecast horizon. iii) *RRMSPE* denotes the ratio of the root mean squared prediction error of the random walk hypothesis to the common factor model. iv) *DMW* denotes the test statistics of Diebold and Mariano (1995) and West (1996). v) *, **, and ** denote rejection of the null hypothesis of equal predictability at the 10%, 5%, and 1% significance levels, respectively. Critical values were obtained from McCracken (2007). vi) Each commodity price is deflated by the U.S. consumer price index to obtain the relative price.

Figure 2.1. Number of factors estimation:
Nominal commodity prices

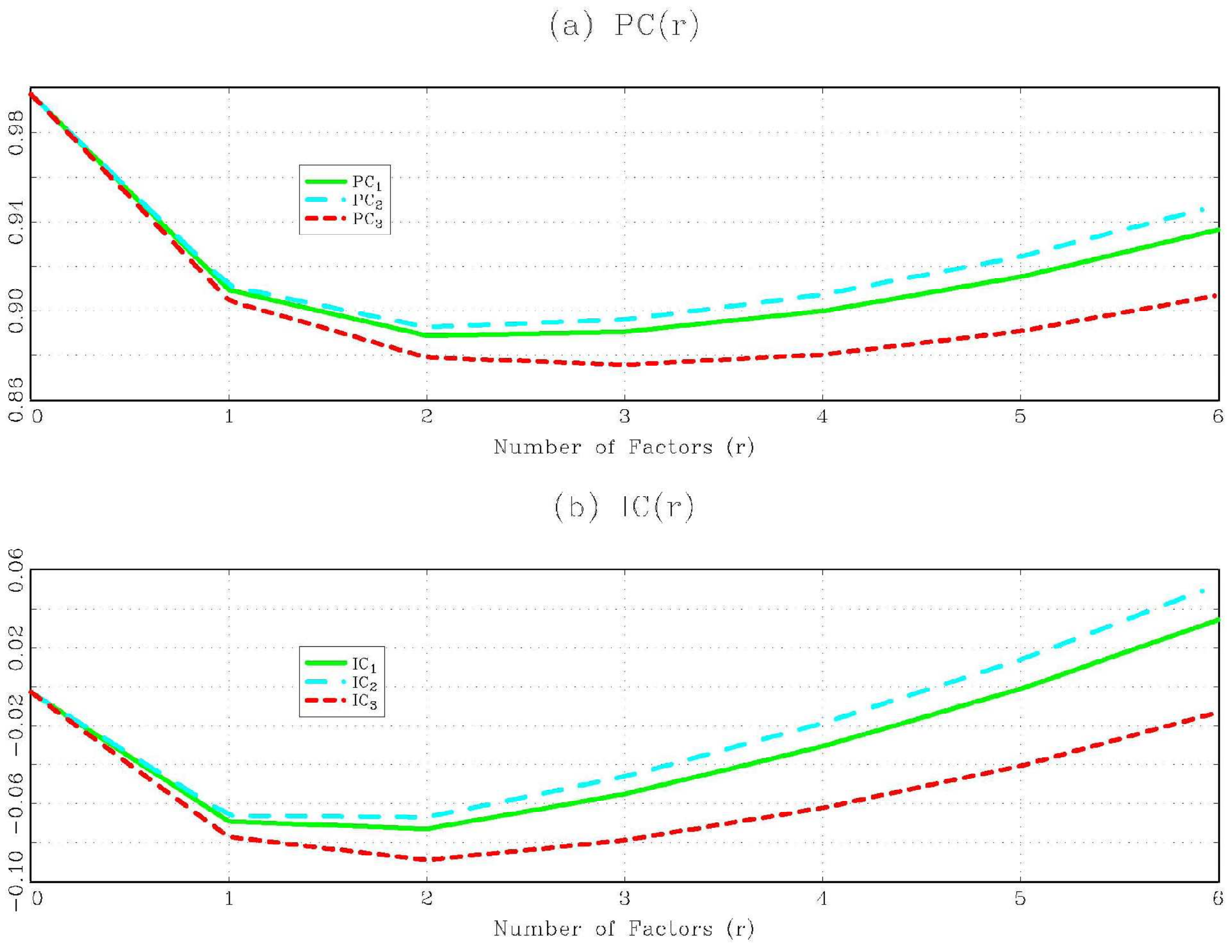


Figure 2.2. Number of factors estimation:
Relative commodity prices

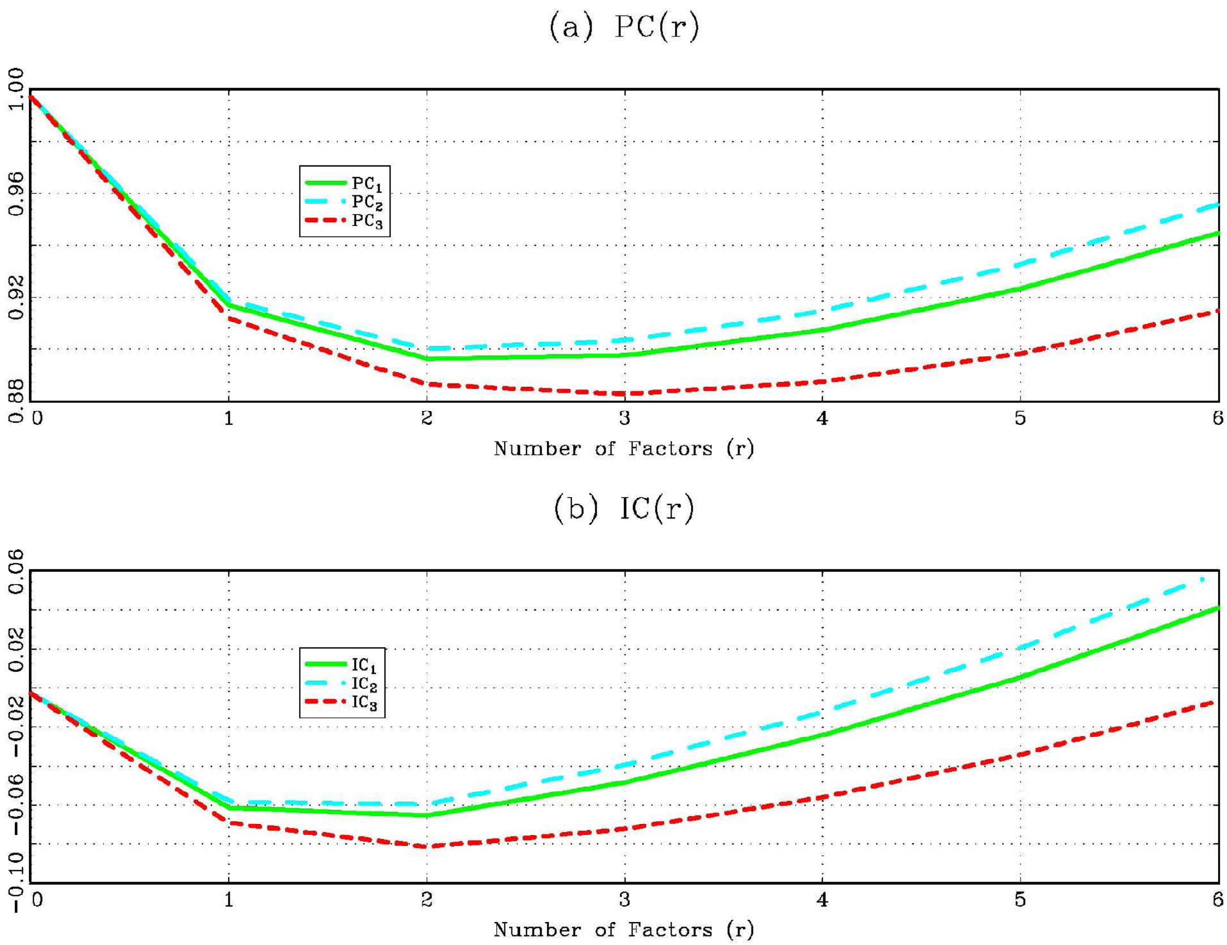


Figure 2.3. Relative importance of common factors:
Nominal commodity prices

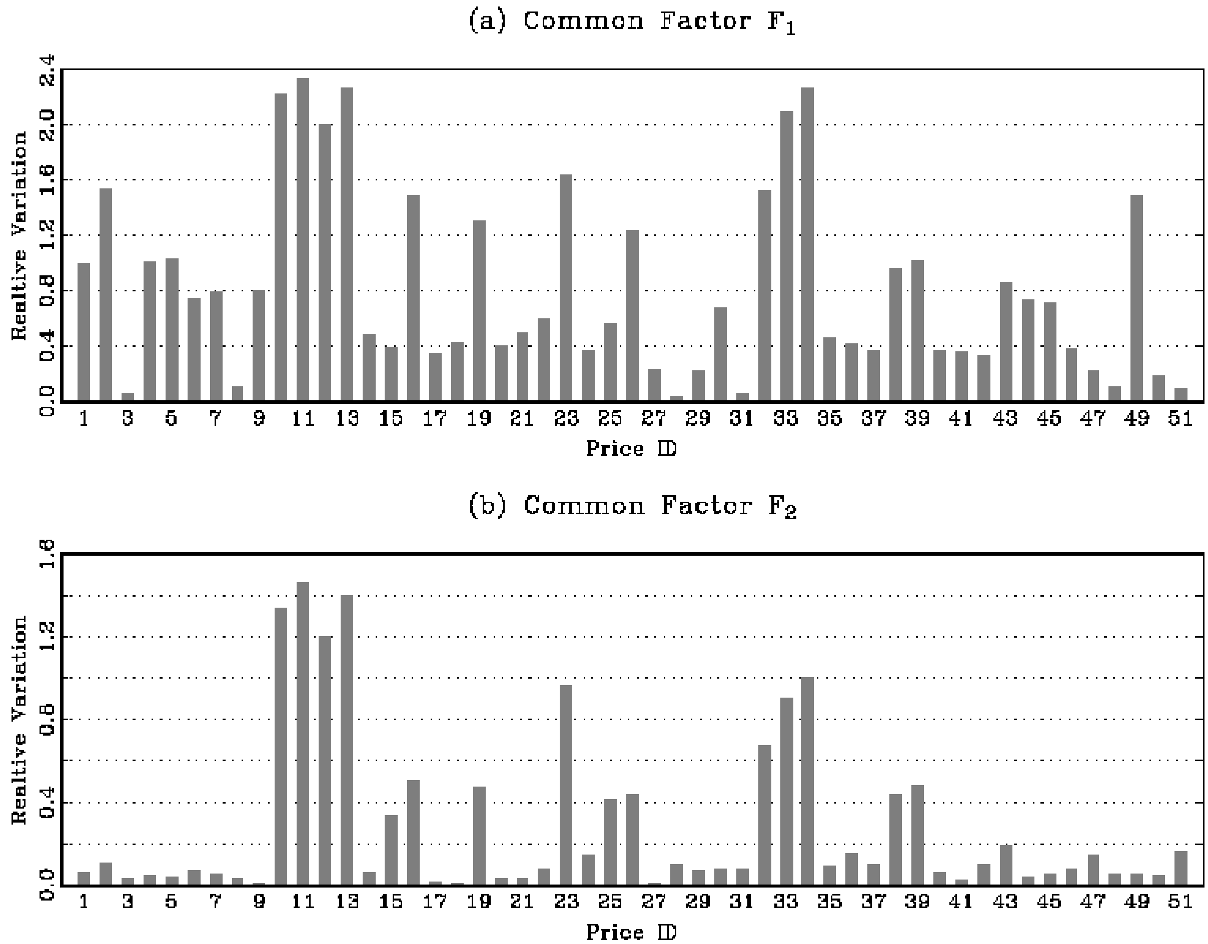
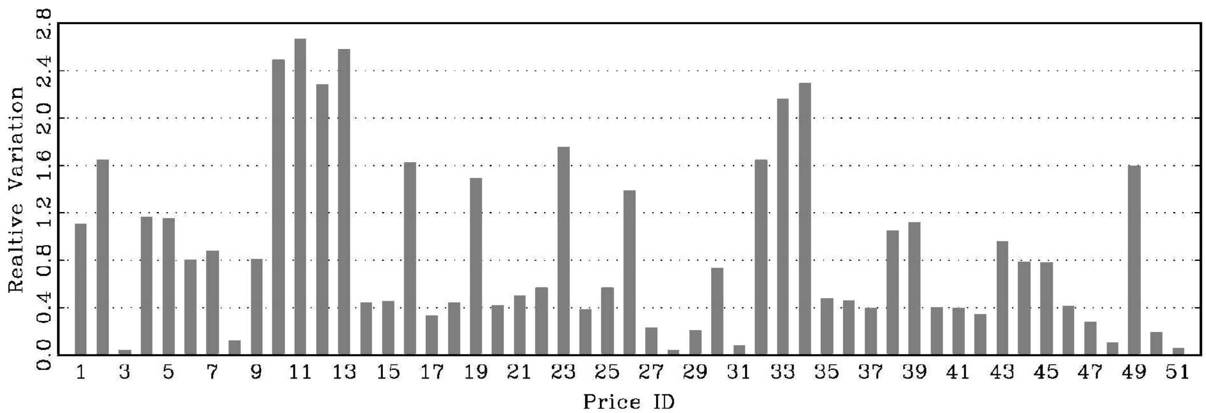


Figure 2.4. Relative importance of common factors:
Relative commodity

prices

(a) Common Factor F_1



(b) Common Factor F_2

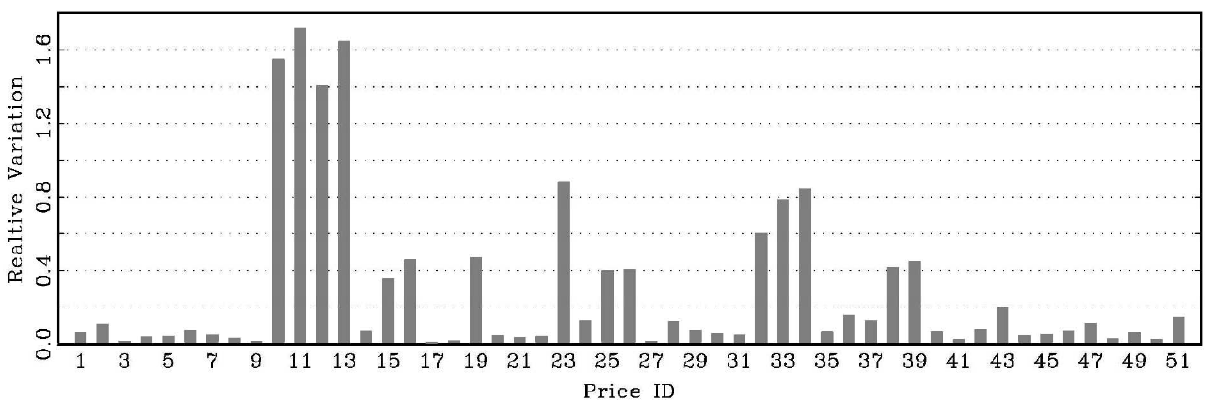


Figure 2.5. Factor loadings estimates:
Nominal commodity prices

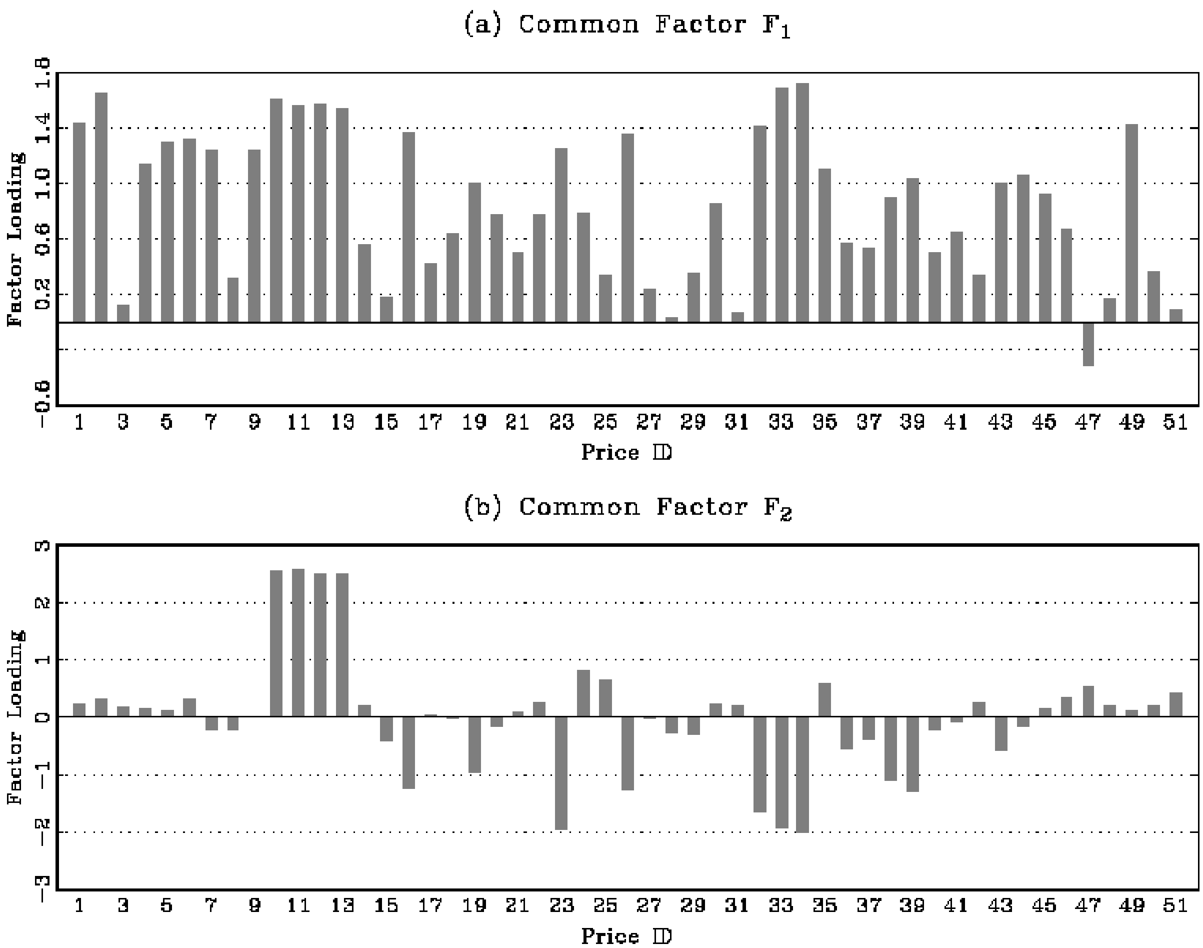


Figure 2.6. Factor loadings estimates:
Relative commodity prices

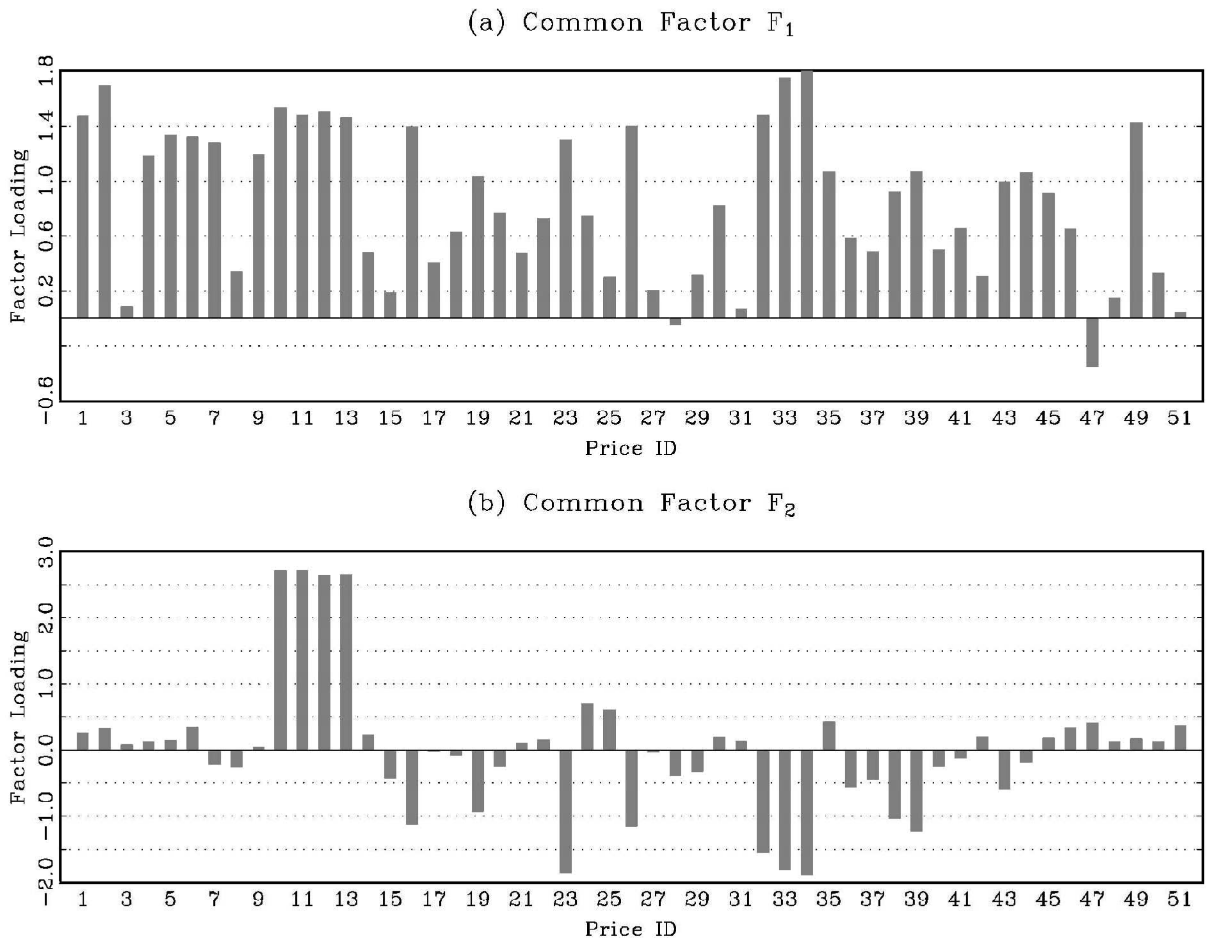


Figure 2.7. U.S. nominal exchange rates vs. common factor estimates

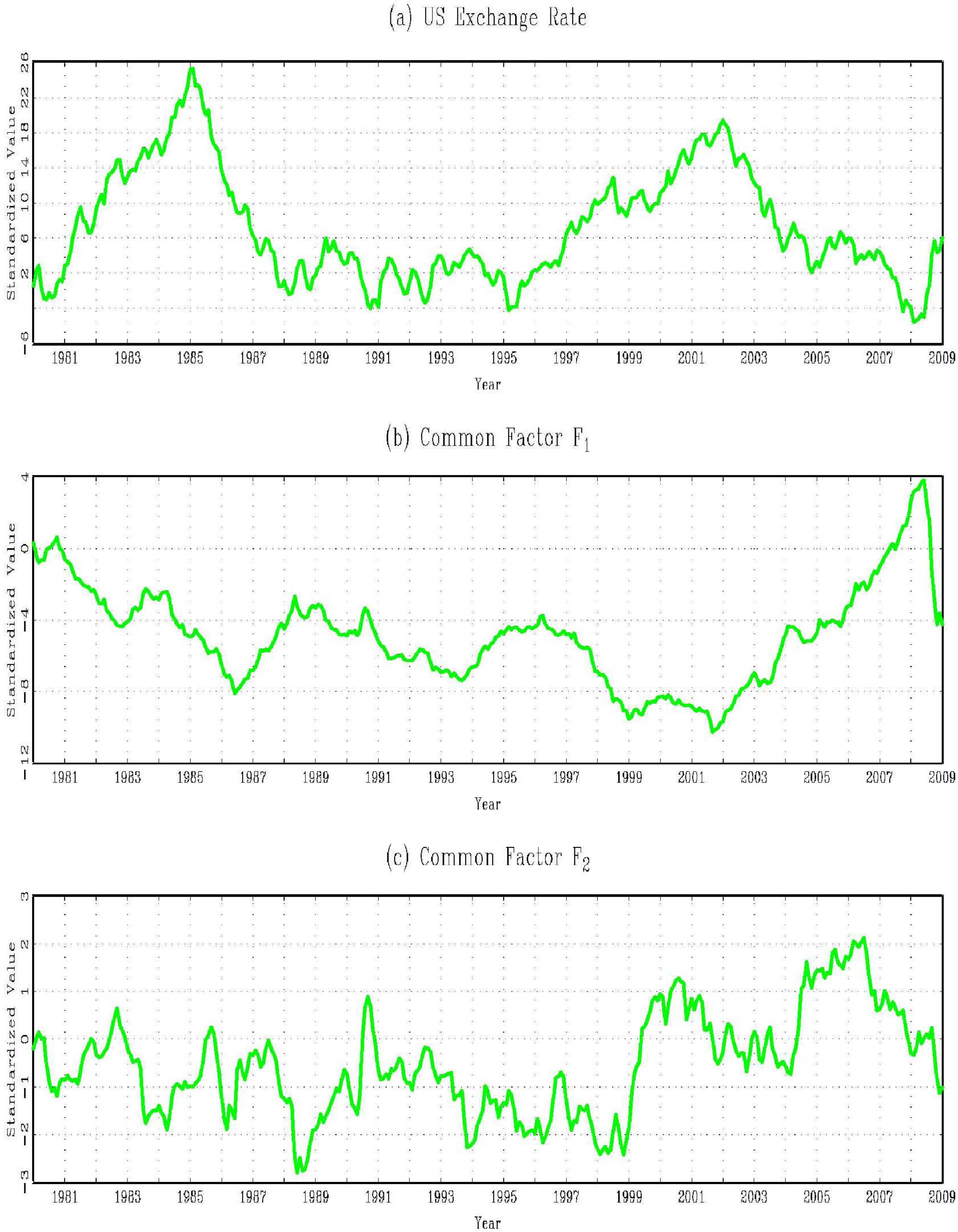
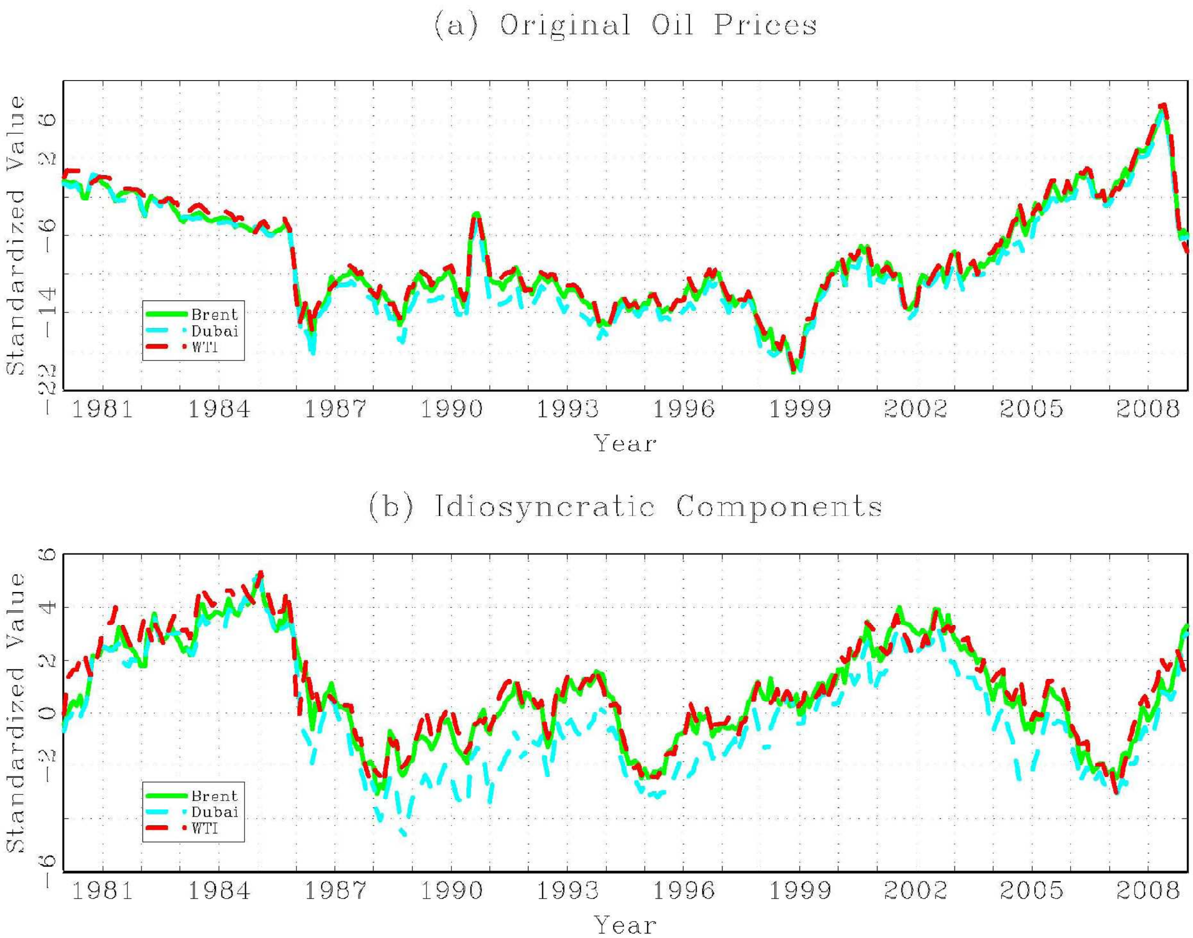


Figure 2.8. Crude oil prices



APPENDIX 3

Tables and Figures for Chapter 3

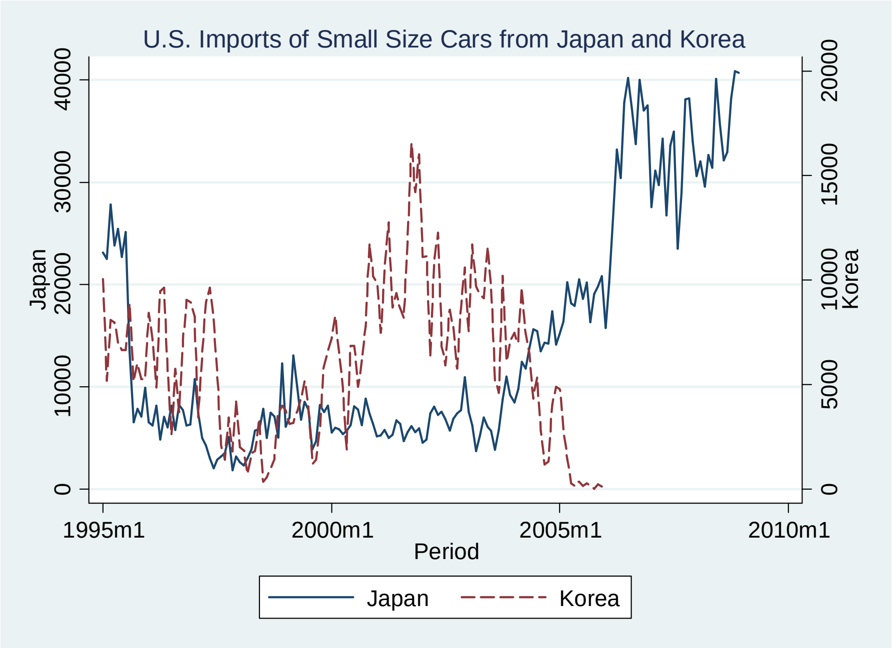
Table 3.1. Import quantity and price data

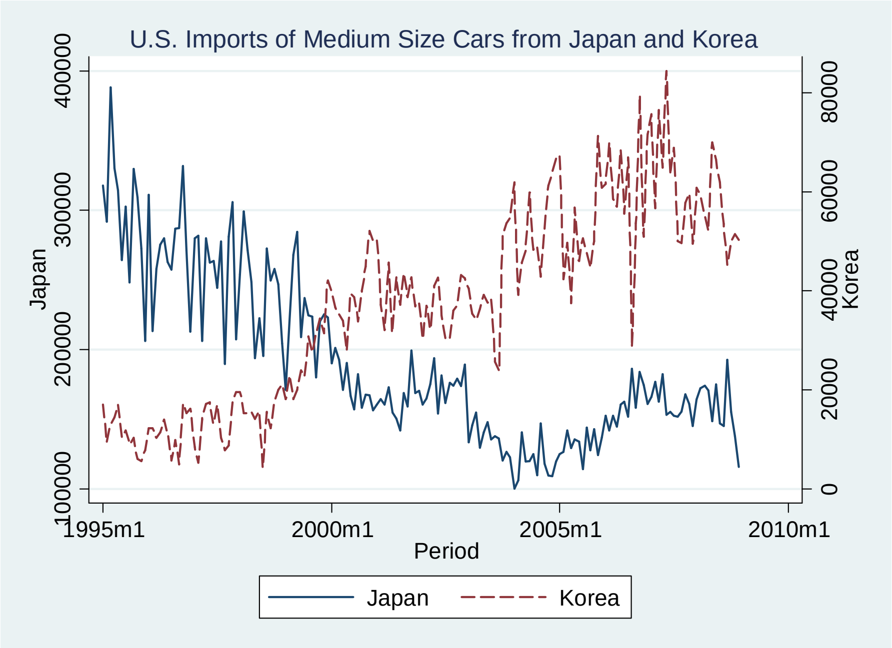
		Small size car import		Medium size car import	
		Japan	Korea	Japan	Korea
1995-2005 (monthly)		c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Import quantity	Average	9,122	6,007	199,235	30,752
	Standard deviation	5,706	3,644	64,342	16,582
Price ($ per unit)	Average	11,154	6,916	8,495	9,646
	Standard deviation	2,050	653	1,932	916
1995-2008 (monthly)					
Import quantity	Average	14,231	-	190,819	36,989
	Standard deviation	11,349	-	59,653	19,603
Price ($ per unit)	Average	11,522	-	8,284	9,082
	Standard deviation	1,960	-	1,784	1,386

Note:
(1) Small size cars: Motor cars and other motor vehicles for transport of persons with spark-ignition internal combustion reciprocating piston engine of a cylinder capacity exceeding 1,000 cc but not exceeding 1,500 cc (HTS 8703.22).
(2) Medium size cars: Motor cars and other motor vehicles for transport of persons with spark-ignition internal combustion reciprocating piston engine of a cylinder capacity exceeding 1,500 cc but not exceeding 3,000 cc (HTS 8703.23).
(3) Small size car imports from Korea discontinued after 2005.
(4) Average price is calculated by dividing total Custom value by quantity of imports.The Customs value is the value of imports as appraised by the U.S. Customs. This value is defined as the price actually paid or payable for merchandise, excluding U.S. import duties, freight, insurance, and other charges.
(5) Price data are deflated using Consumer Price Index.

Data source: United States International Trade Commission (2009) and the Bureau of Labor Statistics (2009).

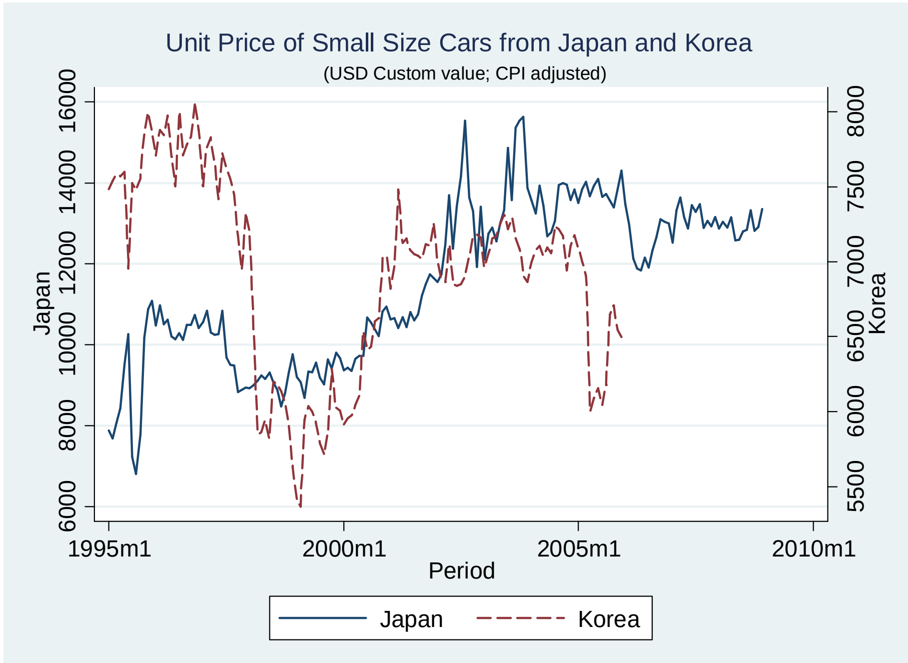
Figure 3.1. Import quantity data series

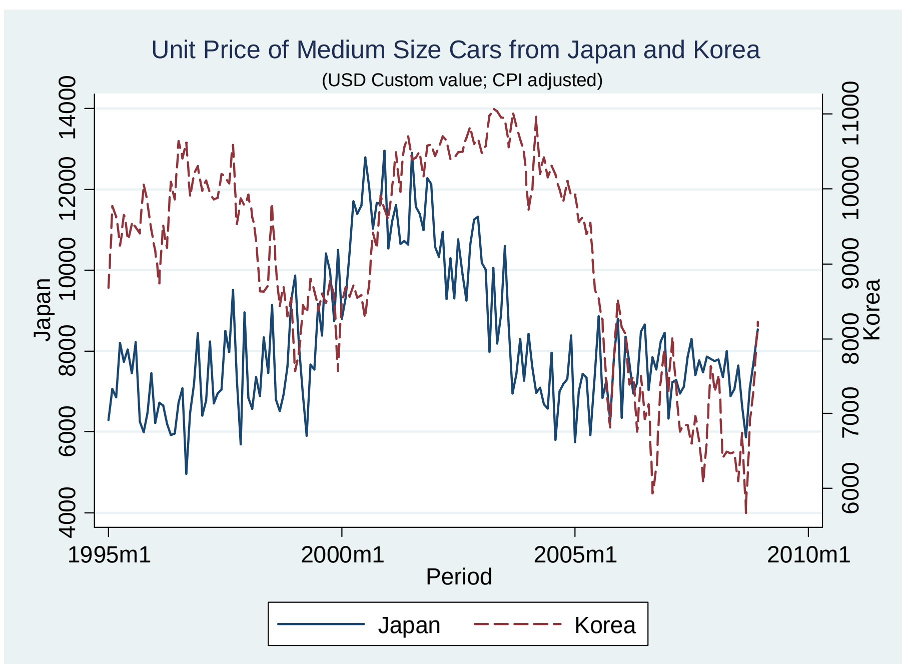




Data source: United States International Trade Commission (2009).

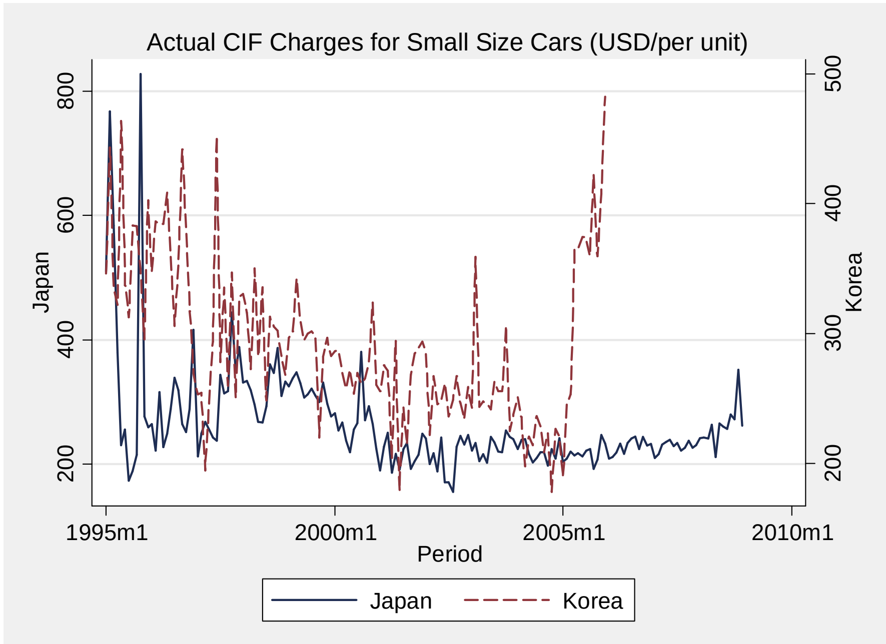
Figure 3.2. Car price data series

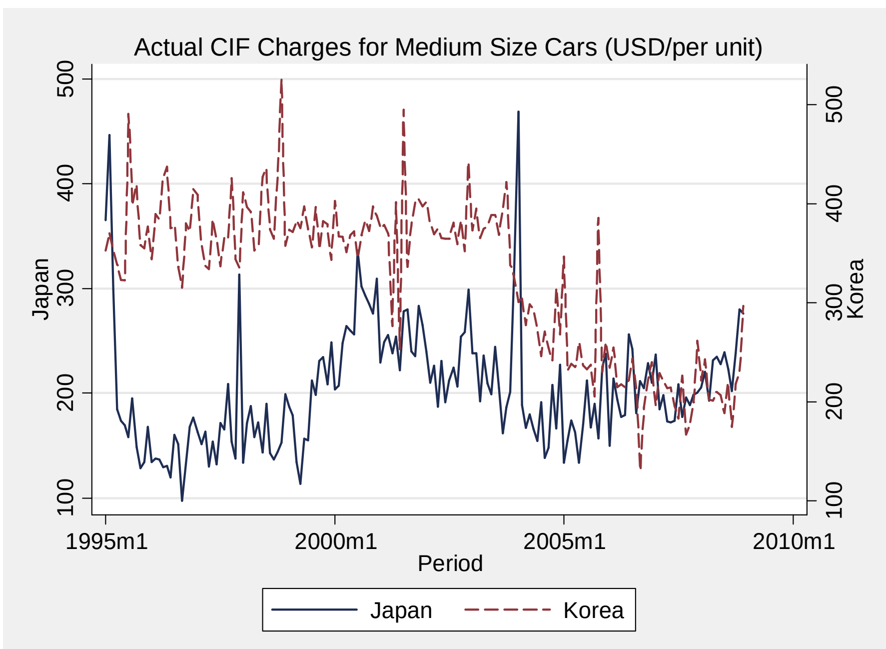




Data source: United States International Trade Commission (2009).

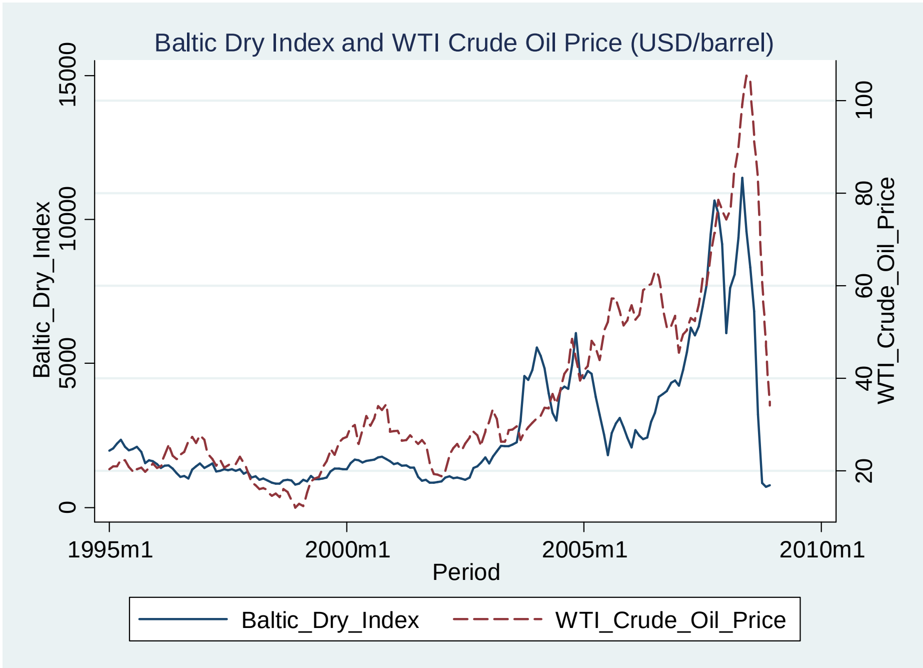
Figure 3.3. Actual shipping cost data series



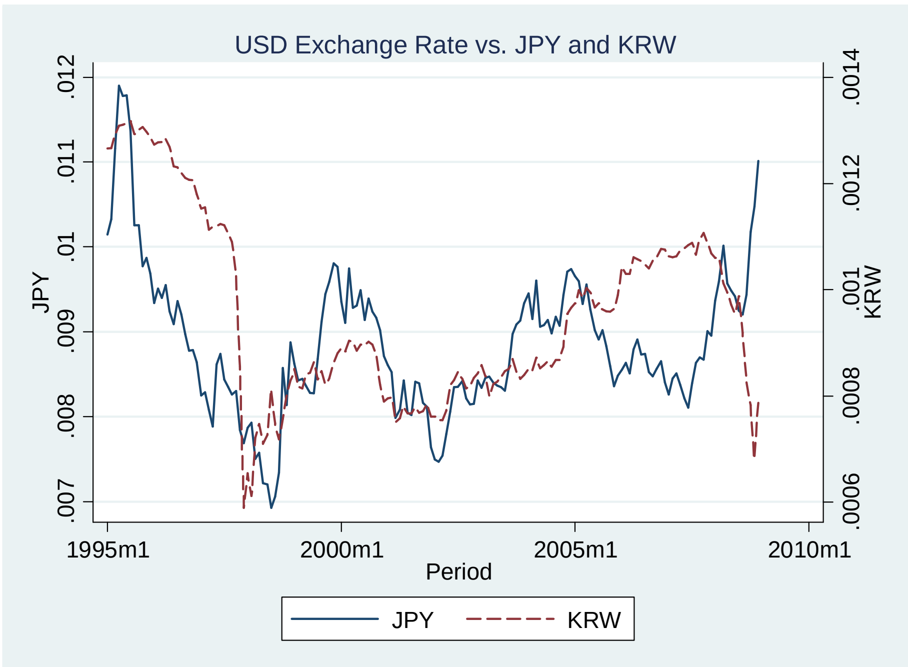


Data source: Calculated based on related data from the United States International Trade Commission (2010).

Figure 3.4. Other data series



Data source: U.S. Bureau of Energy Administration (2009) and FactSet Research Systems (2010).



Data source: Federal Reserve Bank of New York (2009).

Table 3.2. Augmented Dickey-Fuller (ADF) test results

	Variable	ADF test	
		1995m1-2005m12	1995m1-2008m12
Import quantity	c_{sj}^{h}	-3.14*	-2.06
	c_{sk}^{l}	-2.32	
	c_{mj}^{l}	-3.09*	
	c_{mk}^{h}	-2.76	-3.03*
Car price	p_{sj}^{h}	-2.12	-2.54
	p_{sk}^{l}	-2.045	
	p_{mj}^{l}	-4.22*	
	p_{mk}^{h}	-2.78	-2.37
Freight rate	f	-0.50	-1.28
Actual shipping costs	t_{sj}^{h}	-5.85*	-6.58*
(cost, insurance and freight)	t_{sk}^{l}	-5.12*	
	t_{mj}^{l}	-5.35*	
	t_{mk}^{h}	-5.78*	-4.08*
USD/JPY	e_j	-2.25	-1.92
USD/KRW	e_k	-1.93	-1.99

* indicates the result is significant at a 5% level of significance compared to the MacKinnon critical value of 2.89.

Table 3.3. Regressions and cointegration test results (1995m1-2005m12)

(a) Regressions using freight rate

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2005 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price, p_i	0.537*	4.530***	-0.384***	-0.115
	(0.305)	(0.856)	(0.083)	(0.458)
Freight rate, f	0.291*	-0.865***	-0.616***	1.054***
	(0.153)	(0.222)	(0.050)	(0.124)
Exchange rate, e_i	2.954***	3.976***	0.724***	-1.149**
	(0.422)	(0.871)	(0.193)	(0.479)
Intercept	15.78***	-25.33***	23.60***	3.398
	(2.565)	(7.369)	(1.171)	(4.221)
Observation	131	131	131	131
Adjusted R^2	0.415	0.397	0.587	0.413
Breusch-Godfrey LM	66.54	65.77	74.85	82.00
ARCH	13.56	8.11	39.27	14.54
Engle-Granger test EG_τ = -3.8	-4.62	-4.42	-4.21	-3.93

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; standard errors are in parentheses. The freight rate is based on Baltic Dry Index as an increasing function of oil price (with a one-month lag).

(b) Regressions using actual shipping costs (cost, insurance, and freight)

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2005 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price, p_i	1.013***	5.884***	-0.208	0.277
	(0.258)	(1.095)	(0.191)	(0.435)
Shipping costs, t_i	0.142	-1.527***	-0.278*	-1.735***
	(0.179)	(0.484)	(0.146)	(0.251)
Exchange rate, e_i	3.374***	-0.703	0.248	-1.946***
	(0.389)	(0.630)	(0.282)	(0.226)
Intercept	14.713***	-39.95***	16.66***	4.182
	(3.584)	(14.05)	(1.499)	(4.323)
Observation	132	132	132	132
Adjusted R^2	0.399	0.272	0.057	0.463
Breusch-Godfrey LM	69.21	55.19	104.88	55.19
ARCH	20.47	7.63	77.38	7.630
Engle-Granger test EG_τ = -3.8	-4.45	-4.02	-2.94	-5.44

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; standard errors are in parentheses. Actual shipping cost is calculated by deducting the Custom value from the CIF (cost, insurance, and freight) value, and the result is divided by the total quantity of car imports. The Custom value is defined by the USITC (2010) as “the price actually paid or payable for merchandise, excluding U.S. import duties, freight, insurance, and other charges.” The CIF value excludes U.S. import duties.

Table 3.4. Regressions and cointegration test results for higher quality cars, extending the period of observations (1995m1-2008m12)

(a) Regressions using freight rate

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2008 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price, p_i	-0.184	-	-	0.424
	(1.054)			(0.299)
Freight rate, f	1.054***	-	-	0.883***
	(0.096)			(0.092)
Exchange rate, e_i	1.870***	-	-	-1.039***
	(0.418)			(0.388)
Intercept	11.79***	-	-	-0.269
	(2.716)			(3.180)
Observation	167	-	-	167
Adjusted R^2	0.631			0.506
Breusch-Godfrey LM	102.58			100.21
ARCH	26.53			19.58
Engle-Granger test $EG_{\tau} = -3.8$	-4.27	-	-	-4.64

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; standard errors are in parentheses. The freight rate is based on Baltic Dry Index as an increasing function of oil price (with a one-month lag).

(b) Regressions using actual shipping costs (cost, insurance, and freight)

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2008 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price, p_i	2.177***	-	-	0.363
	(0.335)			(0.369)
Shipping costs, t_i	0.293	-	-	-1.457***
	(0.246)			(0.229)
Exchange rate, e_i	3.502***	-	-	-1.468***
	(0.516)			(0.425)
Intercept	3.879	-	-	15.39***
	(4.887)			(2.657)
Observation	168	-	-	168
Adjusted R^2	0.359			0.383
Breusch-Godfrey LM	126.15			99.98
ARCH	74.97			14.58
Engle-Granger test $EG_{\tau} = -3.8$	-3.53	-	-	-4.92

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; standard errors are in parentheses. Actual shipping cost is calculated by deducting the Custom value from the CIF (cost, insurance, and freight) value, and the result is divided by the total quantity of car imports. The Custom value is defined by the USITC (2010) as "the price actually paid or payable for merchandise, excluding U.S. import duties, freight, insurance, and other charges." The CIF value excludes U.S. import duties.

Table 3.5. Stock-Watson dynamic OLS regression results (1995m1-2005m12)

(a) Regressions using freight rate

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2005 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price,p_i	-3.021***	4.97***	-0.486***	-1.916*
	(0.951)	(1.480)	(0.062)	(0.962)
Freight rate,f	2.676***	-0.278	-1.122***	2.628***
	(0.603)	(0.421)	(0.083)	(0.378)
Exchange rate,e_i	-4.395**	-5.635**	-0.288	-5.381***
	(1.824)	(2.602)	(0.717)	(1.357)
Intercept	-3.931	-33.16***	23.37***	8.545
	(10.05)	(11.01)	(3.566)	(7.326)
F value	63.91	64.53	813.3	66.83
Prob > F	0.000	0.000	0.000	0.000

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; Newey-West standard errors are in parentheses. The freight rate is based on Baltic Dry Index as an increasing function of oil price (with a one-month lag).

(b) Regressions using actual shipping costs (cost, insurance, and freight)

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2005 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price,p_i	0.774	4.940***	-0.167	1.591*
	(0.487)	(1.100)	(0.245)	(0.832)
Shipping costs,t_i	0.061	1.122**	-0.445**	-1.750***
	(0.248)	(0.490)	(0.217)	(0.642)
Exchange rate,e_i	4.753***	-12.97***	-1.172**	-1.868
	(0.975)	(3.415)	(0.476)	(2.730)
Intercept	23.859***	-41.19***	10.42***	5.897
	(7.878)	(9.566)	(2.776)	(8.293)
F value	4.03	34.18	15.68	3.46
Prob > F	0.000	0.000	0.000	0.000

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; Newey-West standard errors are in parentheses. Note: c_{mj}^{l} is not cointegrated (see Table 3.3b). Actual shipping cost is calculated by deducting the Custom value from the CIF (cost, insurance, and freight) value, and the result is divided by the total quantity of car imports. The Custom value is defined by the USITC (2010) as "the price actually paid or payable for merchandise, excluding U.S. import duties, freight, insurance, and other charges." The CIF value excludes U.S. import duties.

Table 3.6. Stock-Watson dynamic OLS regression results for higher quality cars, extending the period of observations (1995m1-2008m12)

(a) Regressions using freight rate

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2008 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price, p_i	-1.091***	-	-	1.619***
	(0.417)			(0.480)
Freight rate, f	1.584***	-	-	1.022***
	(0.218)			(0.115)
Exchange rate, e_i	1.189	-	-	-7.577***
	(1.589)			(1.413)
Intercept	12.87	-	-	-12.24**
	(7.756)			(4.919)
F value	24.04	-	-	65.10
Prob > F	0.000			0.000

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; Newey-West standard errors are in parentheses. The freight rate is based on Baltic Dry Index as an increasing function of oil price (with a one-month lag).

(b) Regressions using actual shipping costs (cost, insurance, and freight)

	Small size car import		Medium size car import	
	Japan	Korea	Japan	Korea
1995-2008 (monthly)	c_{sj}^{h}	c_{sk}^{l}	c_{mj}^{l}	c_{mk}^{h}
Car price, p_i	1.654*	-	-	1.175**
	(0.941)			(0.468)
Shipping costs, t_i	-0.413	-	-	-1.551***
	(0.587)			(0.407)
Exchange rate, e_i	10.23***	-	-	-3.481*
	(1.829)			(1.831)
Intercept	44.55***	-	-	8.599**
	(15.63)			(3.352)
F value	3.61	-	-	5.07
Prob > F	0.000			0.000

***,**,* indicate that the coefficient is significant at 1%, 5%, and 10% level, respectively; Newey-West standard errors are in parentheses. Actual shipping cost is calculated by deducting the Custom value from the CIF (cost, insurance, and freight) value, and the result is divided by the total quantity of car imports. The Custom value is defined by the USITC (2010) as "the price actually paid or payable for merchandise, excluding U.S. import duties, freight, insurance, and other charges." The CIF value excludes U.S. import duties.