Trade costs and Deviations from the Law of One Price

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Abstract

The link between relative prices across countries and trade costs requires a theoretical foundation that is absent in the literature to date. Empirical tests on the Law of One Price are generally based on an arbitrage equation, which is of very limited relevance in understanding this link. We develop a new set of arbitrage equations based on a general equilibrium model of arbitrage for commodity markets. The application of the model to the tin trade market yields results that are consistent with our theoretical expectations regarding the link between price differentials and trade costs, after controlling for lag effects and regimes of competition. Generally, we find strong evidence in favor of the LOP between the U.S. and Japan. However, evidence of price discrimination is also found for the Los Angeles market.

Key Words: law of one price, trade costs, arbitrage, commodity markets, market power

JEL Classification Numbers: C21, C22, D40, F10, F41.

1. Introduction

The law of one price (LOP) simply states that in a market free of tariffs, transportation costs, and other trade impediments, identical products sell for the same common-currency price in different countries[1]. Price equalization is often thought to be brought about by goods market arbitrage. According to Rogoff (1996), trade policy barriers and transportation costs drive a wedge between prices in different countries with the size of the wedge depending on the tradability of the good.

Though apparently simple, the LOP raises a number of theoretical and empirical issues. There is an extensive literature aimed at testing the validity of the LOP. Large and persistent deviations from the LOP are well documented in this literature. However, as Anderson and Van Wincoop (2004) point out, this literature still lacks a theoretical foundation necessary to link evidence on relative prices across countries to trade barriers. Most empirical tests are based on an arbitrage equation, known in the literature as Heckscher's commodity points, which has very limited relevance to understanding the link between price differentials and trade costs. They show that in a general equilibrium setup arbitrage will generally lead to price differentials that are much tighter than that implied by the Heckscher's commodity points, when the good is homogeneous and is supplied by exporters located in countries other than those where prices are observed. On the other hand, deviations from the LOP may be persistently larger than the marginal cost of moving the good from one place to the other, mainly due to market imperfections such as monopoly power, imperfect information, and long-term customer-seller relationships, among others[2]. These factors limit the ability of wholesalers to arbitrage price differences[3]. Indeed, in the presence of exclusive national marketing licences, arbitrage costs may be prohibitive for wholesalers.

Therefore, it matters to know the different possible origins (local of production or exporting country) and destinations (local of consumption or importing country), in a general equilibrium setup, to identify who the possible arbitrageurs are and which trade costs are relevant to relative prices between each pair of locations. To our knowledge, none of the empirical studies on the LOP considers a general equilibrium setup and very few have direct measures of trade costs even for bilateral trade[4].

Anderson and Van Wincoop also show that ad-valorem trade costs tend to be quite large, though it is very hard to estimate them with current available data, especially when they include border-related trade barriers and local distributional costs[5]. Therefore, although price differentials between spatially separated markets may often be persistently large, we do not really know how much of these differentials can be explained by trade costs and by other factors[6].

The fact of the matter is that relative prices of similar products sold in different countries have been shown to be systematically related to exchange rate fluctuations between those countries for a variety of goods and countries (Goldberg and Knetter, 1997). For example, Engel and Rogers (1996), comparing retail prices among cities in the U.S. and Canada, show that price dispersion increases with the distance between city pairs, but goes much beyond distance when cities are in different countries. More recently, examining retail price indexes in European countries, Engel and Rogers (2001) find large border effects that are unrelated to trade barriers, but may be explained by exchange rate volatility combined with nominal price stickiness. Parsley and Wei (2001) also find strong evidence of sticky prices in local currency. Furthermore, distance, shipping cost, and exchange rate variability collectively explain a large portion of the border effect between US and Japan[7].

These studies suffer from the already mentioned inappropriate use of Heckscher's commodity points. Since we know the locations where goods are sold, but do not know where they can possibly come from, the relevant

[1] If the LOP is valid, market is said to be integrated. Any perfectly competitive market must be integrated, as price equals marginal cost, but an integrated market is not necessarily perfectly competitive. Segmented markets imply the existence of market power. See Goldberg and Knetter (1997).

[2] In practice, goods are never identical. To the extent that there is some degree of product differentiation, the cost of arbitrage tend to rise further relatively to the marginal cost of moving the good from one place to the other.

[3] Goldberg and Verboven (2001) provide evidence on how small differences in products due to regulatory constraints limited arbitrage in the European car market.

[4] Goodwin et al (1990), Baffes (1991), and Michael, Nobay, and Peel (1994) are exceptions as they explicitly consider shipping costs. Parsley and Wei (2001) use a proxy for shipping and insurance costs based on distance and the cost per unit of distance.

[5] They estimate that ad valorem trade costs are 170 percent for rich countries. Transport costs would account for 21 percent, border-related trade barriers for 44 percent and retail and wholesale distribution costs for 55 percent. These estimates may unintentionally include rents, though rents should not be seen as part of trade costs. Trade costs in developing countries are known to be much larger. See Hummels et al. (2008).

[6] According to Goldberg and Knetter (1997), we do not understand well the sources of international market segmentation or what makes arbitrage costly, enabling substantial price discrimination.

[7] The variability of relative wages is used to control for the non-tradable component of retail prices, but it has little impact or may even increase the border effect.

trade costs for arbitrage are unknown. Given that resale of goods at retail prices tends to be rather costly, arbitrage may be done by producers in other locations, making trade cost between places of consumption irrelevant to explain price differentials. Besides, distance between pairs of consumption locations may not be a good proxy for transport costs, even when constant costs per unit of distance are added[8]. Transport costs depend on several other factors such as the specific good being transported, mode of transportation, volumes being transported, traffic congestion, toll fees, and shippers' market power. The latter depends on the number of shipping firms, the elasticity of demand, product prices, and import tariffs (Hummels et al., 2008). In fact, these studies make no attempt to bring in market power and markups by any of the agents (producers, wholesalers, retailers, and shippers) explicitly into the picture[9]. They are mixed together with border-related trade barriers, local distributional costs, taxes, and regulations that violate with varied degrees the identical goods condition. Finally, as Anderson and Van Wincoop point out, 'there are many factors that contribute to time series variation in relative prices, such as changes in marginal costs, trade costs, taxes, markups, and exchange rates. It is hard to see how information can be extracted about the level of trade costs from evidence on changes in relative prices, especially without the guidance of theory'.

At the retail level it is difficult to satisfy the condition of identical goods[10]. However, using wholesale prices or international trade unit values, many studies have also rejected the LOP for narrowly defined and highly tradable goods. The LOP is rejected for some homogeneous commodities according to early empirical studies (Isard 1977, Richardson 1978, and Kravis and Lipsey 1978). More recently, Thursby et al. (1986), Giovannini (1988), Knetter (1989 and 1993), and Ceglowski (1994), among others, have also provided evidence against the LOP for homogeneous highly disaggregated goods. They all reveal persistent short-run deviations from the LOP[11]. Cointegration-based LOP tests seem to provide more support for the hypothesis. Conceptually, they see price equalization as a long-term equilibrium, but the speed of convergence can be analyzed through a vector error correction model (VECM) as, for instance, in Vataja (2000) and Michael, Nobay, and Peel (1994). On the other hand, cointegration analyses by Ardeni (1989) and Baffes (1991) find mixed results for the same narrowly defined goods. Although most studies recognize the spatial dimension, they fail to take into account the crucial role of time in the process of arbitrage. Since the transportation of goods takes time, arbitrage does not occur directly between current prices, but between forward prices for delivery in a future date. Protopapadakis and Stoll (1983 and 1986) use futures market prices and find support for the LOP.

Since these tests on wholesale prices and trade unit values do not generally control for any other variable, they do not account for any deviation from the LOP. The article by Michael, Nobay, and Peel (1994) is one of the exceptions, because they explicitly consider shipping costs from exporting to importing countries. They show that the LOP cannot be rejected in the long-run and although there are systematic short-run deviations from the LOP, they cannot be predicted ex-ante, hence, no arbitrage opportunity is left unexploited.

Our analysis differs from the previous literature in two main aspects. Firstly, we derive theoretical arbitrage equations, relating international relative prices to international trade costs for highly traded homogeneous commodities, considering different models of competition and lag effects in a general equilibrium framework. These arbitrage equations provide the theoretical foundation necessary to relate deviations from the LOP to trade costs. Secondly, our empirical tests examine the world trade of tin by looking simultaneously at the two main importing countries, the United States and Japan, and the major exporting countries, providing a multi-country framework. C.i.f. and f.o.b. monthly unit values in five different entry ports in the United States allow a direct measure of trade costs and the necessary data to test their relation with price differentials. The border effect between the U.S. and Japan can also be directly estimated. By focusing on the international portion of transportation and on a homogeneous commodity free of policy barriers, we avoid the difficulties associated with measuring border-related trade barriers, local distributional costs, and local markups.

We find that LOP deviations are smaller between the U.S. and Japan than between locations pairs in the United States. Thus no border effect is found for cross-country c.i.f. unit values. Consistent with our theoretical equations, LOP deviations between importing location pairs depend on trade cost differentials from the exporting country, which sells to both locations at the same time, to the importing location pairs. We also find evidence that delivery lags and

[8] Parsley and Wei (2001) estimate the cost per unit of distance between the US and Japan on the basis of the bilateral and proportional difference between c.i.f. and f.o.b. unit values. The intra-national cost per unit of distance is somewhat arbitrarily assumed to be half of the international cost.

[9] Of course local currency pricing means price discrimination and the exercise of market power.

[10] Haskel and Wolf (2001) and Asplund and Friberg (2001), among others, have very creatively managed to maintain the identical goods assumption at the retail level. The former analyzed the retail prices of a set of goods sold by a multinational retailer in different countries, while the latter did the same for goods sold by three Scandinavian duty-free outlets where each good has prices in at least two different currencies. Both studies reject the LOP, though some convergence tends to occur over time.

[11] As time series data on prices tend to be non-stationary, conventional regressions applied in some early studies are invalid.

market power play a significant role in lop deviations between some pairs of locations within the United States. The speed of convergence to the LOP has half-lives that vary from one to four months, the fastest in the literature. Hence, LOP deviations are never persistently large. In contrast, studies that use aggregate price indices, such as the consumer price index, find half-lives of 36 to 60 months for the speed of convergence of real exchange rates. Asplund and Friberg (2001) estimate a half-life of 24 months for the speed of adjustment in the prices of products sold in the Scandinavian duty-free outlets. Studies that focus on individual retail goods across countries find LOP deviations persistency with half-lives in the range of 12 to 24 months. This is the range found in the study by Crucini and Shintani (2008), who analyze a large data set of local currency prices of individual goods and services. They find that the speed of adjustment depends on the degree of tradability of the goods and their inputs, being faster the lower the distribution margin. Half-lives of relative price deviations for automobiles in Europe are estimated in the range of 16 to 19 months by Goldberg and Verboven (2005), while Cumby (1996) finds the half-life of deviations from Big Mac parity to be about 12 months. But our results are best compared with those of Vataja (2000) who estimate that, on average, two thirds of LOP deviations are eliminated within one year for a group of ten internationally traded commodities[12]. In the case of exported tin, he found that from 66% to 83% of LOP deviations are eliminated within one year between Bolivia and Malaysia and Malaysia and Thailand. The equivalent measure in our study for imported tin between the U.S. and Japan ranges from 93% to 100%.

This article is organized as follows. Section two reviews the arbitrage equations often used in the literature to test the LOP and their implicit assumptions. Section three develops a very simple general equilibrium model of arbitrage for homogeneous products. Section four examines the difficulties to find appropriate data to analyze deviations from the LOP, points out the advantages and disadvantages of the available data and describe the methodological strategy applied in the empirical analysis reported in section five. Section six presents some concluding remarks.

2. Arbitrage Equations

Abstracting from trade costs, the price of a homogenous commodity in country i is traditionally (Rogoff 1996) written as:

$$p_i=E_{im}.p^d_m=p_m \qquad (2.1);$$

where E_{im} is the exchange rate of i's currency per unit of m's currency; p^d_m is the price in m's currency; and p_m is the price in country m expressed in i's currency.

Considering t_{mi} as one plus the tax equivalent of the trade costs (transport costs and other trade impediments) between m and i, the following inequalities is frequently referred to in the literature as Heckscher's commodity points:

$p_i \leq t_{mi}.p_m$ (2.2);
and assuming, for simplicity, that trade costs are symmetrical: $p_m \leq t_{mi}.p_i$ (2.3);
which, re-arranging, lead us to the arbitrage equation: $1/t_{mi} \leq p_i/p_m \leq t_{mi}$ (2.4).

Inequality (2.4) has extensively been used as a basis for LOP tests. The threshold autoregression (TAR) models are also based on this arbitrage equation (e.g., Obstfeld and Taylor 1997). It is assumed that relative prices move freely between the arbitrage points, but if they occasionally go outside these limits, they should rapidly be pulled back inside the arbitrage points if the LOP holds true. Though not very clear in the literature, equation (2.4) implicitly considers that arbitrage between markets i and j is carried out by wholesalers or consumers through the resale of the good from one market to the other.

However, as pointed by Anderson and Van Wincoop (2004), if arbitrage is carried out by producers or exporters, then price differentials will generally be smaller than trade costs. Let us review this important point now.

Abstracting from markups, taxes and subsidies in the chain from production to the final user and assuming that the price paid by the buyer is given by the marginal cost of production (c_m) in exporting country m times the trade cost (t_{mi}) from m to i, we can write:

$$p_i=c_m.t_{mi} \qquad (2.5).$$

Country i will then source from producers m=1, 2, …, for which $c_m.t_{mi}$ is the lowest $p_i=\min(c_1.t_{1i}, c_2.t_{2i}, \ldots)$. Arbitrage is done here neither by consumers nor wholesalers, but by producers. If p_i is greater than $c_{mi}.t_{mi}$, it is profitable for the producer in country m to undercut the existing price in country i (Anderson and Van Wincoop 2004). If country i buys from two different countries at the same time, let us say m and n, then it must be that $c_m.t_{mi}=c_n.t_{ni}$.

[12] His annual data is taken from IMF International Financial Statistics for the period 1960-1995 and corresponds to exporter's prices.

If it is optimal for importing country i to source from country z_i and for importing country j to source from country z_j, then

$p_i/p_j= (c_{zi}.t_{izi})/(c_{zj}.t_{jzj}) \leq (c_{zj}.t_{izj})/(c_{zj}.t_{jzj})$; then $p_i/p_j \leq t_{izj}/t_{jzj}$

$p_i/p_j = (c_{zi}.t_{izi})/(c_{zj}.t_{jzj}) \geq (c_{zi}.t_{izi})/(c_{zi}.t_{jzi})$; then $p_i/p_j \geq t_{izi}/t_{jzi}$; then

$$t_{izi}/t_{jzi} \leq p_i/p_j \leq t_{izj}/t_{jzj} \quad (2.6).$$

Clearly, arbitrage equation (2.6) is generally much tighter than (2.4).

Suppose z_i and z_j are the same exporting country (supplier) m, then (2.6) turns into $p_i/p_j=t_{mi}/t_{mj}$. Thus only in this case is relative price completely tied down by trade barriers. And if $t_{mi}=t_{mj}$, then there is no price differential between country i and country j, regardless of the trade cost between i and j! Now if exporting country m supplies importing country i, then $p_i/p_m=t_{mi}/t_{mm}=t_{mi}$. In other words, in this specific case, the price differential between the exporting country and the importing country is equal to the trade cost from m to i. Because time does not play a role here, it is implicitly being assumed that all prices, including trade costs, refer to the same point in time.

On the other hand, price differentials may be larger than trade costs between two importing countries, if markets are not competitive and resale is difficult[13]. Therefore, to properly test the LOP and examine the link between relative prices and trade costs, one has to be explicit about how arbitrage takes place between the different locations and which trade costs are relevant to price differentials.

3. A Simple General Equilibrium Model of Arbitrage

In developing this theoretical model of arbitrage, what we have in mind is a traded homogeneous metal commodity whose world reference price is negotiated at commodity and futures markets such as the LME. Our main goal is to find the implications for the relationship between trade costs and relative price under different sets of basic and extreme conditions, for a highly traded product that is a perfect substitute across exporting countries. Let us start by making a set of assumptions designed to establish a model in which arbitrage could work as perfectly as possible:

(1) The world market is made up of many importing markets in different locations and of some exporting countries with, at least, a few firms in each country.

(2) Exporters decide the total quantity they want to sell on the basis of a reference spot price negotiated at a world commodity market. As a result, supply in each exporting country is fixed at any point in time. This seems to be a reasonable assumption if arbitrage takes place instantly and our concern is with price differentials across different locations at a point in time rather than with changes in world price levels[14].

(3) The demand in each of the importing countries is a continuous function and is negatively related to the commodity price.

(4) Competition is fierce in some large importing markets, so that exporters and shippers take destination price (c.i.f.) as given in these markets.

(5) In these large and competitive markets, exporters also take as given the price of moving the commodity from the exporting country to the entry point in the importing country (c.i.f.-f.o.b.). Although every single shipper takes destination price as given, this assumption does not preclude the possibility that shippers may actually share profits with exporters in these routes or even force them to sell at marginal cost.

As a result of these assumptions, the sole problem for the exporting firm is to distribute its sales among the different importing markets, so as to maximize its instantaneous profits. Let us consider two of these large and competitive importing markets (i and j). Buyers in importing countries always source for the lowest price, while sellers seek for highest prices[15]. As a result, the price (c.i.f.) of this homogeneous commodity at a particular time and destination is expected to be the same, regardless of its origin. Likewise, the price (f.o.b.) of this commodity at a particular time and origin is expected to be the same independently of its destination. In other words, neither importers nor exporters can price discriminate. International trade costs, including shipping times (see Anderson and Van Wincoop 2004), are given and are known to both importers and exporters. Relative prices between different locations at a point in time are here constrained not by a bilateral relationship between one importer and one exporter, but by a general equilibrium version of

[13] This encompasses a number of more specific assumptions such as: products not being identical; shipping time differences, introducing risk and uncertainties; impossible resale due to long-term customer-seller relationships; among others.

[14] Indeed, it is not unusual for an exporter of tin to auction a certain quantity of a commodity to be shipped at a given date to a particular location. We shall have more to say on this in the next sections.

[15] Profits are here obtained by exporters and/or shippers. These are short run profits determined by demand raising prices above average costs, rather than by markups derived from sellers' market power.

arbitrage in which there are several potential exporters/origins and importers/destinations. We would like to focus first on an instantaneous equilibrium situation, in which each destination is supplied by at least two different exporting countries, and at least one exporting country supplies both destinations at the same time with equal shipping times.

Proposition I: if the above conditions hold, then the following non-arbitrage condition must also hold for any exporter m that sells to both locations i and j at the same time:

$\log(p^*_j)-\log(p^*_i) = \log(t_{mj})-\log(t_{mi})$ (3.1),

where p^*_i and p^*_j are the prices for which demand equals supply in importing countries i and j; t_{mi} and t_{mj} are one plus the trade costs, including shipping firms' profits if any, from m to i and j, respectively[16]. In other words, proportional price differentials between two importing countries are determined by the exporter (arbitrageur) that sells to both locations, and are equal to the proportional difference between the two destinations of the trade costs from this exporting country.

Proof: Let us define $p^*_{mi}=c_m.(1+r_{mi}).t_{mi}$, where c_m is the marginal cost of the product produced in exporting country m and r_{mi} is the rate of profit of exporters when exporting from m to i. Analogously, $p^*_{mj}=c_m.(1+r_{mj}).t_{mj}$[17].

If m is a country that exports to both import destinations i and j, then exporter's profits per unit (Π_m) to both destinations must be equal: $c_m.r_{mj}=\Pi_{mj}=c_m.r_{mi}=\Pi_{mi}$ => $c_m.(1+r_{mj})=c_m.(1+r_{mi})$ => $p^*_{mj}/p^*_{mi}= t_{mj}/t_{mi}$ => $\log(p^*_{mj})-\log(p^*_{mi})=\log(t_{mj})-\log(t_{mi})$. Since importers pay the lowest price for the product that is expected to arrive at each destination at the same time, regardless of country of origin, then $p^*_{mi}=p^*_i$ and $p^*_{mj}=p^*_j$. Therefore, $\log(p^*_j)-\log(p^*_i)=\log(t_{mj})-\log(t_{mi})$.

Corollary I: Under the above conditions, the proportional trade cost difference from the same exporter to destinations i and j must be smaller (larger) than the proportional trade cost difference between the same destinations from any exporter that sells only to the destination where price is the lower (higher) of the two at the same point in time. That is:

For $p^*_i<p^*_j$

$\log(t_{mj})-\log(t_{mi})< \log(t_{zj})-\log(t_{zi})$ (3.2),

for any exporter z which only exports to i, and

$\log(t_{mj})-\log(t_{mi})> \log(t_{zj})-\log(t_{zi})$ (3.3),

for any exporter z which only exports to j.

Proof: If $p^*_i<p^*_j$, we can write from (3.1) that $t_{mi}<t_{mj}$. Any exporting country z which exports to country i, where the price is lower, but not to j, will have larger profits per unit of exports when exporting to i than to j ($\Pi_{zi}> \Pi_{zj}$): $c_z.(1+r_{zi})>c_z.(1+r_{zj})$. But as $p^*_i=c_z.(1+r_{zi}).t_{zi}$ and $c_z.(1+r_{zi}).t_{zj}>p^*_j$ then $p^*_j/p^*_i= t_{mj}/t_{mi}< c_z.(1+r_{zi}).t_{zj}/c_z.(1+r_{zi}).t_{zi}$ => $t_{mj}/t_{mi}<t_{zj}/t_{zi}$ => $\log(t_{mj})-\log(t_{mi})<\log(t_{zj})-\log(t_{zi})$. If $p^*_i<p^*_j$, but z exports to j (higher price) and not to i, then $\log(t_{mi})-\log(t_{mj})<\log(t_{zi})-\log(t_{zj})$ or $\log(t_{mj})-\log(t_{mi})>\log(t_{zj})-\log(t_{zi})$.

Corollary II: Under the above conditions, if the trade cost of an exporting country z to destination i, where price is lower, is at least as high or higher than to destination j, then z will only export to j, where price is higher, and the trade cost difference between the two destinations of this exporter will be either zero or will have the opposite sign to the trade cost differentials of the arbitrageur to the same destinations.

Proof: $p^*_i<p^*_j$ and, for exporting country z, $t_{zi}\geq t_{zj}$. From $p^*_i<p^*_j$ it follows that $c_z.(1+r_{zi}).t_{zi}<c_z.(1+r_{zj}).t_{zj}$ if z could export to i and j; but since $t_{zi}\geq t_{zj}$ then $c_z.r_{zi}=\Pi_{zi}<c_z.r_{zj}=\Pi_{zj}$. Therefore, country z will only supply importing country j. Since $t_{zi}\geq t_{zj}$ then $\log(t_{zj})-\log(t_{zi})\leq 0$, but since $p^*_i<p^*_j$ then $\log(t_{mj})-\log(t_{mi})>0$.

It is worth noting that price differentials between exporting and importing countries must be equal to the trade costs between them, as in Heckscher's commodity points. In contrast, price differentials between importing countries are equal to the trade cost differential of shipping the product from exporting countries that are indifferent as to whether they export to one or to the other importing country.

[16] Because it takes time for the goods to move from one place to the other, there is a lag between the price observed in m and the prices observed in i and j. However, as shipping time from m to i and j is assumed to be the same, all prices and trade costs are assumed to be negotiated and settled at the same time prior to the embarkation of the product.

[17] Note that p^*_i is equivalent to the c.i.f. price and $c_m.(1+r_{mi})$ is equivalent to the f.o.b. price. Trade costs may include profits and if total profits go to shippers, then $r_{mi}=0$ and total profits are included in t_{mi}.

Let us now relax some of the assumptions made at the start of this section. Let us now assume that an exporting firm at country m is a price taker when exporting to destination i, but is a monopoly when exporting to destination j. In other words, the cost (cif) of the product shipped to j by the second most efficient exporter is prohibitive. We initially assume that the exporting firm in m can extract all profits so that trade costs from m to i and from m to j include no profits. Therefore, given t_{mi}, the exporting firm will maximize profits where p^*_i/t_{mi} (the fob price) equals its marginal cost mc(x). Profits must be non-negative for m to be an exporter to i. The exporting firm total supply (x*) is thus determined by this profit maximization. When exporting to j, the monopolist firm will then equate its marginal revenue to its marginal cost. That is:

$$\max_{x_{mj}} (p^*_j/t_{mj}).x_{mj} - (p^*_i/t_{mi}).x_{mj}$$

The first order condition will give: $(p^*_j/t_{mj}).(1-1/|\varepsilon|)=(p^*_i/t_{mi})$. Therefore, the proportional price difference between the two destinations will now be:

$$\log(p^*_j)-\log(p^*_i) = \log(t_{mj})-\log(t_{mi})-\log(1-1/|\varepsilon|) \qquad (3.4).$$

Therefore, monopoly in one destination market widens price differentials compared to other destinations, independently of trade costs differentials. Note that the same result will be obtained if shipping firms take the fob price in m as given, but while they are confronted with a perfectly elastic demand curve in i, there is a monopoly service from m to j[18]. However, in this case, f.o.b. price differentials will be equal to zero, while it will be different from zero if the exporter is a monopolist.

Clearly, if the cost (cif) of the product shipped to j by the second most efficient exporter from country z is lower than the monopoly price of the good from m, the monopolist can practice limit pricing in the Bertrand-Nash equilibrium, thereby driving the rival from the market. We will then have:

$$\log(t_{mj})-\log(t_{mi})\leq\log(p^*_j)-\log(p^*_i)\leq\log(t_{mj})-\log(t_{mi})-\log(1-1/|\varepsilon|). \qquad (3.5).$$

That is, price differentials will be in between the competitive case and the full monopolist case. Market power implies that only one exporter will sell in the market at a point in time.

Let us now consider some lag effects. Suppose that, giving shipping times from an exporter m to destinations i and j, price and trade cost have to be settled one period earlier for buyers in i than in j, for the good to arrive at the same time:

$$p^*_{jm}(t)= p_m(t-1).t_{mj}(t-1)$$

$$p^*_{im}(t)= p_m(t-2).t_{mi}(t-2),$$

where p_m is the f.o.b. price at m. Thus,

$$\log[p^*_j(t)]-\log[p^*_i(t)]=\log[t_{mj}(t-1)]-\log[t_{mi}(t-2)]+\log[p_m(t-1)]-\log[p_m(t-2)] \qquad (3.6).$$

Current price differentials between two destinations that buy from the same exporter depend now on the changes over time in f.o.b. prices and in trade costs of the exporter.

Consider a second exporter z that only ships to destination j. If it takes longer for z's product to arrive at j than m's product, importers at j will buy from z if they expect, based on information at the time Ω(t-k), when they buy from z, the c.i.f. price from m at (t-1) to be higher than the c.i.f. price from z at (t-k):

$$E\,[p_m(t-1).t_{mj}(t-1)|\,\Omega(t-k)] > p_z(t-k).t_{zj}(t-k) \qquad (3.7)$$

But if they do not know how c.i.f. prices will change from (t-k) to (t-1), they might split their purchases between the two countries. The result is that the current c.i.f. price differentials at one destination from different exporting countries will no longer be zero.

$$\log[p^*_{jm}(t)]-\log[p^*_{jz}(t)]=\log[t_{mj}(t-1)]-\log[t_{zj}(t-k)]+\log[p_m(t-1)]-\log[p_z(t-k)] \qquad (3.8).$$

[18] This is probably a more realistic assumption. For a theoretical and empirical analysis of market power in international shipping, see Hummels et al. (2008).

4. Data and Methodological Strategy

Non-alloyed tin appears to be an appropriate product to test the LOP and our model of arbitrage for homogeneous commodities. Alongside copper, tin was the first metal to be traded on the London Metal Exchange (LME). LME tin futures contract specification requires, among other things, that the product is of 99.85% minimum purity. Although most LME contracts do not include the physical sale of the product, it seems that almost all non-alloyed tin imported in large markets satisfy this LME standard[19]. Therefore, non-alloyed tin is likely to be one of the most homogenous tradable commodities, though small quality differences remain[20].

The international market for non-alloyed tin has two large importing countries: the United States and Japan[21]. Tin has been imported by the U.S. and Japan free of tariff and non-tariff barriers, regardless of country of origin, since at least 1989. The largest net exporters of tin are Indonesia, Peru, Malaysia, Thailand, China, Bolivia, and Brazil[22].

The U.S. imports largely from Peru, Bolivia, Brazil, China and Indonesia. Japan, on the other hand, imports mostly from Indonesia, China, Malaysia, and Thailand. Singapore, Malaysia, Thailand and Hong Kong import large volumes from Indonesia and seem to operate as a hub for Indonesia's exports. They also seem to refine part of Indonesia's tin exports, since the government in Indonesia struggles to enforce the prohibition of exports of non-alloyed tin below the standard set up by the LME[23].

Data on tin imports from the U.S. and Japan used in this empirical work are monthly values in their respective currencies and quantities in weight units. Import unit values (c.i.f. and f.o.b.) on a monthly basis are available in digital form in the U.S. by port of destination and country of origin as from January 1989[24]. Import unit values (c.i.f.) on a monthly basis are also available in digital form in Japan by country of origin for the same period[25]. Import unit values were calculated as the ratio of import values to quantities in U.S. dollars per ton and yens per ton. Japan import unit values were converted from yens into dollars using the current monthly average exchange rate[26]. All time series were transformed in logs. LME spot prices are also available in digital form on a daily basis from July 1989 to date[27].

The lack of appropriate data is one of the major difficulties to test the validity of the LOP and to examine the underlying arbitrage mechanism. Monthly unit values are far from the ideal proxy for international trade prices. Ideally, one would like to have data on forward contract prices between exporters and importers in different locations. Alternatively, one would like to have data on unit values for every ship from each exporter to each importer with dates of departures and arrivals. Unfortunately, data on these prices and unit values are not publicly available.

According to tin traders, forward prices are contracted on the basis of LME daily spot prices. Usually, prices are set equal to the average of fifteen to twenty previous working days of LME spot prices, some five days before the product is embarked, plus a premium that varies according to demand and supply short term conditions, quantity embarked, and minor differences in the quality of the product. Because import unit values are available only on a monthly basis, they represent a weighted average of LME daily spot prices with some lags. The problem is that we do not know the weights, which depend on unknown daily quantities and dates of shipments within each month to each destination. A comparison between the maximum and minimum values within each calendar month of a lagged moving average of daily LME spot prices and import unit values in Japan and the U.S. appears to confirm this tin pricing rule.

In order to examine the arbitrage mechanism, it is very useful to take advantage of the information on tin imports by exporting country and port of destination available for the United States. This disaggregated data allows comparisons of import unit values at each country of origin (f.o.b.) to different ports of destination and at each port of destination (c.i.f.) from different countries of origin. This also permits a critical examination of the data and a fine identification of unit value outliers, improving the data quality. The main disadvantage of the data, as mentioned before,

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[19] The quality of Indonesian exports appears to be occasionally lower, as it will be seen in the next section.

[20] Tin is mainly used for welding and as an input for electronic products and components. For some specific purposes of the electronic industry, the quality required may be higher than the LME standard.

[21] They accounted for 53% of the eight largest net importers of tin in the world in 2005 and 2006 according to Comtrade database: http://comtrade.un.org/db/.

[22] Singapore, the UK and the Netherlands are net importers of non-alloyed tin, although they are also significant exporters. See Comtrade.

[23] Illegal mining and smelting in Indonesia has often been reported. See, for instance, American Metal Market, March 25, 2005.

[24] Non-alloyed tin is defined at the 10-digit level of the U.S. Harmonized Trade System: HTS 8001100000. See http://dataweb.usitc.gov/.

[25] Non-alloyed tin is defined at the 9-digit level of the Japanese Trade System: 800110000. See http://www.customs.go.jp/toukei/info/index_e.htm.

[26] The average current monthly exchange rates (yens/dollar) are from the IFS database. The introduction of different lags was tested without any visible improvement in the results.

[27] See http://www.lme.co.uk/ and http://www.econstats.com/home.htm.

is the fact that we are working with ex-post averages rather than prices. Furthermore, disaggregation has a cost. Time series have many missing values, as exporters do not export every month to each destination port, and unit values are sometimes based on low imported quantities. Volatility of unit values tends to increase with the lack of regularity in shipments, both in terms of frequency as well as size of shipments. The main methodological implication is that with many missing values or few observations, econometric time series analysis tends to become inapplicable and statistics less reliable.

Our empirical analysis is basically divided in three steps. First we look into c.i.f. unit values by different exporters at a single destination. Our objective here is to test if products from different exporters are identical, there are no lag effects and competition prevails, so that we cannot reject the hypothesis that their c.i.f. unit values are the same. Second we examine c.i.f. unit values of spatially separated markets. Our main goal is to test the LOP. Our third step is to analyze the link of unit value differentials to trade costs and the presence of market power.

In the first two steps we apply two statistical procedures. Whenever possible, we use time series econometrics, applying the ADF test to see if individual series of unit values are non-stationary and Johansen cointegration test to see if we find a cointegrating vector for each pair of unit value series. If there is a cointegrating vector, we test for zero-mean reversion, applying the Vector Error Correction Model (VECM) with and without the constant term and restricting the coefficients to (1,-1). The speed of adjustment is also tested by applying an impulse and response test (Cholesky d.f. adjusted one standard deviation innovations) and measuring it by its half-life in months. When it is not possible to apply this procedure due to too many missing values in the series, we run hypothesis tests for the mean and median of the c.i.f unit value differential series. If the observations are normally distributed, we apply a two-tailed t-test to the zero-mean hypothesis, otherwise we apply a non-parametric sign test to check the zero-median hypothesis.

In our third step we also apply two different statistical approaches. First we test the zero-mean or zero-median hypotheses for the series of f.o.b. unit value deviations from each exporter to different pairs of spatially separated destinations. Rejection of these hypotheses for identical products being shipped simultaneously suggests the possibility of market power being exercised by exporters. If these hypotheses cannot be rejected, but the c.i.f. unit value deviations of different pairs of spatially separated destinations from a single exporter are significantly larger than trade costs differentials, there is a possibility of market power being exercised by shipping companies in specific routes. Finally, we directly test the following arbitrage equation for two pairs of spatially separated destinations in the United States:

$$\text{lopdev} = c(1) + c(2)\ \text{tradecost} + c(3)\ \text{fobuvch} + \varepsilon \tag{4.1}$$

where lopdev is the unit value differential in log between two different destination markets at a particular month, considering all exporters to these destinations: $\text{lopdev} \equiv \log(p^*_j)_t - \log(p^*_i)_t$; tradecost is the trade cost differential in log from one exporter (arbitrageur) to the same pair of destinations at the same month: $\text{tradecost} \equiv \log(t_{mj})_t - \log(t_{mi})_t$; and fobuvch is the change in the f.o.b. unit value in log of the exporter (arbitrageur) with respect to the previous month: $\text{fobuvch} \equiv \log(p_m)_t - \log(p_m)_{t-1}$. Since our time series of c.i.f. unit value differentials, trade costs differentials and change in the f.o.b. unit values have proven to be stationary, we regress these variables, applying the method of least squares. On the basis of our theoretical arbitrage equations, we expect c(1) to be zero and c(2) to be positive and close to one if neither exporters nor shippers have market power in both destinations, while c(3) should be different from zero if there are shipping lag effects.

5. Results

Let us start by analyzing U.S. imports of tin by port of destination. Baltimore has been the main port of entry into the U.S., accounting for 61% of total tin import volume (in tones) in the nineteen years period from 1989 to 2007. New York comes in second with 22%, while New Orleans, Los Angeles and San Francisco account for 6%, 4% and 3%, respectively. The remaining 4% is spread out over twenty six other customs districts in the United States. In fact, New York used to be the main port of entry for tin in the first five years of this time series, but it tended to decline in importance over time, accounting for only 2.6% of U.S. tin imports in the five years period from 2003 to 2007. Baltimore, on the other hand, accounted for 81% of total imports in this more recent period. Given its size, we shall consider Baltimore the main port of reference in the U.S. tin import market and shall compare it to the other districts with a view to analyzing the arbitrage process in this country. Before that, however, we shall first examine the unit value differentials between different exporting countries at each of the main ports of destinations in the U.S. and in Japan as a whole. The objective is to test the LOP at a point in space and gather information about our data. That should provide some parameters that will be useful to the analysis of spatially separated markets.

5.1 C.i.f. Unit Value Differentials at one location from different countries of origin

5.1.1 U.S. main ports

Table (1) shows that c.i.f. unit value differentials from the two main exporting countries at each month to the five largest destination ports in the U.S. have median and mean very close to zero and small standard deviation[28]. The hypothesis that the observations are normally distributed cannot be rejected for the destination ports of Baltimore, New Orleans, Los Angeles, and San Francisco. The hypothesis of zero-mean cannot be rejected by two-tailed *t*-tests in any of these four destinations (See Appendix (I)). The hypothesis that the median of unit value differentials in New York is zero cannot be rejected by a large sample non-parametric sign test[29].

Therefore, we cannot reject the hypothesis that c.i.f. unit values for the two main exporting countries are the same at each destination. The observed errors, as measured by the mean of the absolute values (absdev mean), average around 4% in the East Coast (Baltimore and New York), 5.0% in the Gulf (New Orleans) and 5.6% in the West Coast (San Francisco and Los Angeles) of the United States. Bearing in mind that we are considering just the two main exporters at each destination port, we take this measurement error as largely accounted for the fact that monthly unit values are being used instead of daily actual (forward) prices. Daily forward price differentials at a single destination from different exporters are expected to depend on buyers' imperfect information and differences between exporters with regards to shipping time, quantities shipped and quality of product. But these differences tend to be smaller when only the two main exporters are being considered. On the other hand, they tend to be larger, the larger the change in price over time. Therefore, the smaller and less frequent the shipments from the main exporters to each destination are, the larger monthly averages of daily forward price differentials will tend to be. Considering these means of the absolute values in each destination, larger deviations are never persistent since they never last for more than three months in any of the series[30]. These error measurements will be useful as a reference in comparison with LOP deviations between different destinations. In any case, we may conclude that LOP deviations for the two main exporters are neither large nor persistent at any particular port of destination.

Considering the two main exporting countries to Baltimore in each month, we have found that both time series of c.i.f. unit values for the first and second main exporters were non-stationary, while the difference in logs was stationary (see ADF Tests in Appendix III). Therefore, the Armington elasticity of substitution is infinite and tin is a perfect substitute by country of origin in the long run. Trace test indicates one cointegrating vector at the 1% level between the two series of unit values[31]. The standard error indicates that the intercept is not significant. Without a constant, the coefficients of the endogenous variables are close to (1,-1). Imposing this restriction on the coefficients and applying the VECM, with no constant, we cannot reject the hypothesis that the coefficients are indeed (1,-1). Therefore, the unit values of the two main exporters reverse to a zero-mean. This may be seen as evidence for the homogeneity of imported tin in Baltimore by country of origin during this period. Impulse and response tests based on VECM indicate that the speed of adjustment is very high, since the response of the leading exporter to the follower and of the follower to the leader have both half-lives of just one month. Thus, monthly c.i.f. deviations are not persistent. In point of fact, c.i.f. deviations are never greater than 5% or smaller than -5% for more than two consecutive months over this whole 19-year period. Due to the number of missing observations, it is not possible to replicate the above time series analysis for the data of New York, New Orleans, Los Angeles and San Francisco.

5.1.2 Japan

Assuming Japan is small enough to be spatially considered a single destination, the time series of c.i.f. unit values of each of the largest four exporters to this import market is also examined. Table (2) provides some basic descriptive statistics for the series of the c.i.f. unit value differentials in Japan by countries of origin. We note that the standard deviations and the mean of the absolute value of the standard deviations for the two main exporters (first column) are even lower than in Baltimore. This seems to reflect greater regularity in terms of frequency and quantities shipped by each of the main exporting countries. Medians and means are close to zero, except perhaps for Indonesia-Thailand and Indonesia-Malaysia. Normality tests reject the hypothesis that c.i.f. unit value differentials have a normal distribution. Applying a large sample non-parametric sign test, we can not reject the hypothesis of zero-median for all the series, except when Indonesia is one of the countries. Apparently c.i.f. unit values of tin from Indonesia are lower than those from China, Thailand and Malaysia.

Again, ADF tests reveal that the time series of c.i.f. unit values of all these exporters are non-stationary, while the differentials of all pairs of exporting countries are stationary. We can not reject the hypothesis that there is one

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[28] Although the period of analysis is from January 1989 to December 2007 (228 observations), the number of observations is smaller because in some months either there is no import at the destination port or imports come from one exporting country only.

[29] See Appendix (II). Nonparametric Sign Tests and Zero-Median Hypothesis Tests. For a description of the test see, for instance, Wonnacott and Wonnacott (1972), p.397-402.

[30] In the case of San Francisco, in addition to the fact that there are many months without imports from two different countries, imports from Indonesia in some months were excluded from the sample as outliers.

[31] Cointegration is always analyzed through Johansen tests in this paper. See Appendix IV: Cointegration Tests, VECM and Impulse and Response Tests.

cointegrating equation between each pair of variables at the 1% level. Applying VECM and imposing the restriction (1,-1) on the coefficients, the standard error always indicates that the intercept is not significant and the restriction cannot be rejected without constant, except for two pairs: Indonesia-Thailand and Indonesia-Malaysia. Impulse and response tests indicate that half-lives are one month for all pairs, but those two. Therefore, we may conclude that unit value differentials tend to zero and the speed of adjustment is very fast indeed. Considering that these exporters supply Japan with great regularity and that they are all in the same region, the two exceptions suggest that tin exported by Indonesia might, at times, not fulfill the condition of identical product, exporting a lower quality tin.

In the case of Indonesia and Malaysia, c.i.f. unit values cointegrate with a constant at the 1% level. The VECM indicates that the constant term is significant, but the restriction (1,-1) is rejected with and without the intercept. Impulse and response tests estimate half-lives of one month. Therefore, the speed of adjustment is very fast, but LOP deviations do not reverse to zero. In the case of Indonesia and Thailand, c.i.f. unit values cointegrate with a constant at the 1% level. The restriction (1,-1) is rejected without the intercept, but cannot be rejected with it and the constant becomes significant. Impulse and response tests indicate that half-life of Indonesia's response to a shock in Thailand's unit value is 3 months, but Thailand response is only 1 month. Therefore, given the sign of the constant term, we conclude that Indonesia's unit values tend to be smaller than Thailand's and the response of the latter to the former is much faster than the opposite.

Unit value deviations between the two main exporting countries at each month, the leader and the follower, are never greater than 5% or smaller than 5% for two consecutive months over this whole 19-year period. One-sided deviations larger than 5% for more than two consecutive months are rare events for the six pairs of main exporters to Japan, but they occurred for three consecutive months seven times, for five consecutive months three times and once for six consecutive months. Thus, although unit value deviations in Japan are small, as standard deviations indicate, somewhat larger and a bit more persistent deviations do occur between pairs of unit values of the main exporters[32].

5.2 U.S.A.-JAPAN

Now that we have a good idea of the basic statistics of the monthly series of c.i.f. unit value deviations at the main ports of entry in the U.S. and in Japan, we can compare them with c.i.f. unit value deviations between the U.S. and Japan. Table (3) reports some basic descriptive statistics of these deviations between the U.S. and Japan. Compared to Tables (1) and (2), medians and means are as close to zero, and standard deviations or the absolute deviation means are as small, except, perhaps, when the West of the U.S. is one of locations.

The individual series for each location are all non-stationary, while the series of LOP deviations between each pair of locations are stationary. The series of the U.S. cointegrate with the series of Japan at the 1% level, but the constant is not significant both with and without the restriction that the coefficients are (1,-1). The hypothesis that the coefficients are (1,-1) has a probability of 79% without a constant. Therefore, we can not reject that LOP deviations between the U.S. and Japan reverse to a zero-mean. Impulse and response tests indicate that the U.S. response to a one standard deviation shock in Japan has a 3 months half-life, while the half-life response of Japan to the U.S. is just 2 months. The speed of adjustment is slightly less fast than when we were dealing with different exporters to just one destination. Nevertheless, the fact that LOP deviations are small (small standard deviations) and not persistent provides strong evidence in favor of the LOP.

If we apply the same econometric procedure to the series of c.i.f. unit values, but considering only the five main exporters to the U.S. and the four main exporters to Japan, we get basically the same results, except that the restriction (1,-1) on the coefficients is rejected without constant and cannot be rejected with constant. The intercept is close to zero (0.005), but is significantly different from zero. LOP deviations reverse to this mean, whose sign indicates that the c.i.f. unit values are lower in the U.S. than in Japan. This result suggests that fringe competitors tend to raise the average of unit values in the U.S. and the dispersion of LOP deviations between the U.S. and Japan as shown in Table (3). The half-life responses of the U.S. to Japan and vice-versa go down to just two and one month, respectively, when only the main exporters are being considered, suggesting that fringe competitors react more slowly to changes in prices; i.e., arbitrage is less efficient.

Bearing in mind the long distance and significant trade costs between the East and the West sides of the U.S., we have also applied the same econometric procedure to test LOP deviations between Baltimore and Japan and between Los Angeles and Japan[33]. We find that LOP deviations reverse to zero-mean between Baltimore and Japan, since the constant is not significant and we cannot reject the hypothesis that the coefficients of the cointegrating equation is (1,-1) without the constant term. Japan responds very quickly to price shocks in Baltimore (half-life=1 month), but Baltimore reacts more slowly to shocks in Japan (half-life=4 months). On the other hand, LOP deviations between Los Angeles and Japan reverse to a mean (0.015) whose sign indicates that unit values in LA tend to be higher on average than in Japan.

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[32] Bearing in mind that we have a sample of 1368 observations (6 pairs times 228 months), those larger than 5% one-sided deviations for three, five and six consecutive months show frequencies of only 1.5%, 1.1% and 0.4%, respectively.

[33] The series consider all exporters to each location. Two missing values in the series of unit values of Baltimore were filled with the unit values of New York, while five of Los Angeles were filled with those of San Francisco.

But the response of Los Angeles to shocks in Japan has a half-life of just 1 month, while the response of Japan to Los Angeles has a half-life of 2 months. Therefore, although unit values tend to be higher in LA, the arbitrage mechanism appears to be more efficient between these locations than between Baltimore and Japan whose unit values tend to be the same. It is noteworthy that standard deviation is much higher for the LOP deviations between Los Angeles and Japan than between Baltimore and Japan. This reflects the fact that c.i.f. unit values in LA are much more volatile than in Japan and Baltimore. The small size of the LA market is likely to play a major role in the volatility of unit values there. The large size of the Baltimore import market might be part of why Japan responds so fast to changes in unit values in Baltimore, while the much smaller size and lower trade costs might explain why LA responds so fast to Japan. But a more complete explanation has to wait until we examine the arbitrage process and the trade costs on the basis of our arbitrage models later in this section.

LOP deviations between Japan and the East or West of the U.S. were also tested by the same econometric procedure, but considering only the 5 main exporters to the U.S. (US 5) and the four to Japan (JAP 4) and merging the markets of Baltimore and New York, called the US East, and the markets of Los Angeles and San Francisco, the US West. The results differ because the constant term between the East and Japan is now significant, though very small (0.007), and the constant term between the West and Japan is now not significant. Unit values are marginally smaller in the East than in Japan, which are about the same as in the West. Japan continues to respond faster to price shocks in the East (half-life=1 month) than the other way around (half-life=2 months). The West and Japan respond to each other very quickly with half-life equal to one month. Testing US East and US West directly confirms that the constant term is significant, so that the series reverse to a negative mean equal to 0.012, indicating that unit values in the East are significantly smaller than in the West. The two areas of the U.S. respond to each other with half-lives of just one month.

5.3 U.S. Pairs of Ports

We have also tested LOP deviations between pairs of the main ports of entry in the United States, considering all exporters to each destinations. Table (4) shows the basic descriptive statistics. Standards deviations tend to be larger between the main ports in the U.S. than between the U.S. and Japan, as shown in Table (3), probably reflecting the greater volatility of unit values in the smaller import market destinations in the United States. We find that the time series of unit values to each destination are all non-stationary, but the time series of the LOP deviations between each pair of port destinations are all stationary.

Applying the Johansen cointegration test to the unit values by each pair of port destination, we cannot reject the hypothesis that there is one cointegrating equation at the 1% level for all pairs. Applying VECM and impulse and response tests, we find that the speed of adjustment is again very rapid, with half-lives of one month, for all pairs of destinations in both directions, except for San Francisco to New York with a half-life of two months. When we impose the restriction (1,-1) on the coefficients of the VECM, we find that we cannot reject this restriction for all pairs of unit values and the constant is not significantly different from zero for the following pairs: Baltimore-New Orleans, Baltimore-New York, Baltimore-San Francisco, New York-New Orleans, New York-San Francisco and New Orleans-San Francisco. In these cases, therefore, unit values reverse very rapidly to the zero-mean, suggesting that arbitrage is effective and we cannot reject the hypothesis that the arbitrageurs' trade costs are approximately the same for these pairs of destinations. Note, however, that whenever Los Angeles is one of the destinations, the constant term is always significant and the sign of the intercept indicates that unit values at the port in LA are larger than elsewhere. Although LOP deviations may be somewhat large at times among the main ports, the fact that mean reversion is very fast indicates that large LOP deviations are not persistent, providing evidence to the LOP.

Another approach is to construct confidence intervals for the means and medians of the LOP deviations between U.S. pairs of ports and test the hypothesis that they are zero. None of the series of LOP deviations follows a normal distribution, thus we must apply a non-parametric test. The last three lines of Table (4) report the results of a large sample non-parametric sign test. The test confirms that unit values in LA are greater than in BA, NO, SF and NY (5%). It also confirms that the following equalities cannot be rejected: BA=NY (10%), NY=NO (10%), NO=SF (5%), BA=NO (10%), BA=SF (1%), LA=NY (1%) and NY=SF (1%)[34].

5.4 Relating unit value differentials between pairs of destinations to trade costs and market power

U.S. import data allow us to investigate further the arbitrage mechanism among the main ports of entry in that country and to try to relate unit value differentials to trade costs and market power. We take pairs of the major destination ports and calculate the LOP deviations for f.o.b. values [$\log(p_{mj})-\log(p_{mi})$] and for c.i.f. values [$\log(p^*_{jm})-\log(p^*_{im})$] for each exporting country m that exports simultaneously (arrivals in the same month) to destinations i and j. Country m is what we called an arbitrageur in section 3 of this text. Because of the significant number of missing values in these series, we can not apply time series econometrics. Alternatively, we construct confidence intervals for the median of LOP deviations based again on a nonparametric sign test.

[34] In parentheses are the levels of significance.

Table (5) shows some basic descriptive statistics and the confidence intervals for the LOP deviations of f.o.b. values. It is worth recalling that these are f.o.b. values reported by the importing country (the U.S.). Hence, they are f.o.b. values of shipments that arrived within the same month at the two destinations, but may have departed and had their prices negotiated at different moments in time. If we cannot reject the hypothesis that the medians of these f.o.b. LOP deviations are zero, then we cannot reject the hypothesis that exporters from a particular country cannot persistently discriminate different importing markets. Furthermore, if f.o.b. LOP deviations were exactly zero, then c.i.f. LOP deviations would be equal to the difference between the trade costs (including shippers' profits) to the two destinations from the arbitrageur.

We can see in Table (5) that we cannot reject the hypothesis that medians are equal to zero for LOP deviations measured at f.o.b. values for most exporting (arbitrageurs) countries and pair of destinations, even at the 10% significance level (SL). Therefore, for almost all exporting countries and pairs of destinations, exporters are unable to price discriminate a particular market in a systematic way. The two exceptions are for China, when exporting to LA in comparison with Baltimore and San Francisco. Chinese f.o.b. unit values are higher when exports are to Los Angeles than when they are to Baltimore or San Francisco.

Table (6) shows some basic descriptive statistics and confidence intervals for the LOP deviations of c.i.f. unit values between pairs of destinations by exporting country. Again for the vast majority of cases, we cannot reject the hypothesis that LOP deviations have zero medians. All four exceptions include Los Angeles as one of the destinations. This is absolutely consistent with our previous results. c.i.f. unit values are higher in Los Angeles than in Baltimore for imports coming from China and Peru. They are also higher in Los Angeles than in San Francisco for imports coming from China and are higher in New York than in Los Angeles when coming from Indonesia.

We already know that LOP deviations reverse to a zero-mean, when all exporting countries are considered, except when Los Angeles is one of the destinations. Table (5) and (6) reveal that medians of LOP deviations of arbitrageurs are not zero precisely when Los Angeles is one of the destinations. Thus, our main task is to try to explain why c.i.f. unit values are higher in Los Angeles and approximately the same in the other main destinations.

We can note from Table (5) that Peru and Brazil were the main arbitrageurs between Baltimore and New Orleans, while Peru was the sole arbitrageur between New Orleans and Los Angeles. Peru exported to Baltimore and New Orleans simultaneously in 117 months, while Brazil did so in 57 months. Peru exported to New Orleans and Los Angeles simultaneously in 71 months. In Baltimore, Peru, Bolivia and Brazil were the main exporters of tin with 29%, 19% and 17% of the import volume in the period from 1989 to 2007. Brazil was the most frequent exporter, present in 224 months out of 228 months. Peru was the most frequent (121 months) and the largest exporter to New Orleans, accounting for 75% of the total volume imported at this destination in the period. Brazil came second, shipping in 57 months and accounting for 13% of the total imported volume.

We already know that equality of c.i.f. unit values between the two main exporters at each month to Baltimore or to New Orleans cannot be rejected. We cannot reject the hypothesis that c.i.f. unit value differentials are zero between Brazil and all the other exporting countries or between Peru and all the other exporting countries both in Baltimore and in New Orleans. We cannot reject the hypothesis that f.o.b. and c.i.f. LOP deviations are zero between these two destinations for imports from Peru and Brazil. Therefore, Peru and Brazil were competitive in both markets and able to check each other's price and take any opportunity for arbitrage. Thus, we find no evidence of market power by exporters or shippers from Brazil and Peru in either Baltimore or New Orleans.

Our model predicts that c.i.f. price differentials between these two destinations should be equal to the ad valorem trade cost differentials of the arbitrageurs between these two locations. In fact, ad valorem trade costs are relatively small in proportion to import unit values (log tm= log c.i.f. – log f.o.b.)[35]. As a result, trade cost differentials between pairs of destinations in the U.S. from any particular exporting country (log tmi – log tmj) are almost always negligible. Brazil's and Peru's trade cost differentials between Baltimore and New Orleans have median, mean and standard deviations equal to (0.0002;-0.0019;0.0232) and (-0.0016;-0.0018;0.0048), respectively. Thus, LOP deviations tend to be approximately zero between Baltimore and New Orleans and are never persistently large.

Baltimore-New Orleans

We have tested equation (4.1) for Baltimore and New Orleans. When there was more than one arbitrageur in a particular month, we calculated the trade cost differential variable as a weighted average, considering as weights the quantities exported from each arbitrageur to these destinations. The changes in f.o.b. unit values were also calculated as a weighted average, considering as weights the quantities exported by the arbitrageurs to each destination. We find that the constant term is never significant. This makes more robust our result that there is no evidence of market power being exercised by exporters from the main arbitrageurs in Baltimore and New Orleans. The regression includes 170 observations and yields:

[35] Ad valorem trade costs vary from one to three percent and ad valorem trade cost differentials are around 0.3% when Los Angeles is not one of the destinations.

lopdev = 0.849962 tradecost - 0.285483 fobuvch

(0.250679)	(0.059767)	(standard error)
[3.390645]	[-4.776603]	[t-statistic]

Since we have observed that the dispersion of unit value differentials increases as exported quantities by country decline, we have attempted to correct any heteroscedasticity by dividing both the dependent and the independent variables by the exponential of $\log(5x10^6)-\log(q_m)$, where q_m is the total quantity exported by the arbitrageur to the two destinations and $5x10^6$ is the monthly quantity average exported by the main countries to Baltimore in the 19-year period. This yields a clear improvement in the precision of the regression coefficients, as shown by the decline in the standard errors:

lopdev = 1.130010 tradecost - 0.202990 fobuvch

(0.215705)	(0.041228)	(standard error)
[5.238689]	[-4.923544]	[t-statistic]

The positive and close to one coefficient of the trade cost differential variable is consistent with our theoretical expectation for competitive markets.

Indonesia (41%; 157 months), China (27%; 102 months) and Peru (25%; 78 months) were the largest and most frequent exporters to Los Angeles. Bolivia was a marginal supplier, shipping in only 12 months, whereas Brazil never exported to Los Angeles. Thus, Los Angeles has largely been supplied by Asian exporters. But Peru is the main connection between New Orleans and Los Angeles, supplying these two markets within the same month for 71 months. Exporters of Peru do not seem to be able to systematically price discriminate these two markets, as indicated by the descriptive statistics and the nonparametric confidence interval for the median of f.o.b. LOP deviations in Table (5). On the other hand, Chinese f.o.b. unit values are often higher for exports to LA than to San Francisco or Baltimore. China's c.i.f. unit values are also higher in LA compared to San Francisco or Baltimore. Peru's c.i.f. unit values are higher in Los Angeles than in Baltimore too[36].

There are three possible non-excludent explanations for China's f.o.b. unit values to LA being higher than to other destinations: differences in shipping time from China to these locations lead to prices being settled at different moments in time; the quality of Chinese tin is higher when exported to Los Angeles than to the other destinations; and Chinese exporters have market power and can price discriminate for a significant period of time. The first hypothesis is hard to believe since the routes and distances from China to Los Angeles and San Francisco are about the same, and China's trade costs to Los Angeles are on average lower than to San Francisco and to Baltimore, eliminating the possibility of LA being a time-consuming and expensive port. The second hypothesis is not consistent with the data either. Why would China sell such a superior product only to LA? Furthermore, when we test the c.i.f. unit value deviations between China and Peru in Los Angeles, we find that the nonparametric confidence interval at the 10% level of significance includes the zero median. Thus, we cannot reject the hypothesis that their unit values are the same at this destination. This also implies that tin from China and Peru is a homogeneous product in LA. Therefore, we can not reject our third hypothesis that Chinese exporters can at times exercise market power in LA. Recall that Peru's c.i.f. unit values are higher in Los Angeles than in Baltimore. Since Peru's f.o.b. unit values are the same to these destinations, Peru's c.i.f. unit values are higher in Los Angeles, because trade costs are in fact higher to that destination than to Baltimore[37]. Therefore, in the case of Peru, it is the shipping service between Peru and Los Angeles that is able to exert market power and benefit from higher prices in Los Angeles.

It is interesting to note that Indonesia and China very often exported simultaneously to Los Angeles and Baltimore. Given that trade costs from these exporters to Los Angeles are much lower than to Baltimore, one could expect prices in Baltimore to be higher than in Los Angeles. The opposite is true, because Peru, Bolivia and Brazil lead exports to Baltimore, cutting prices there. If these markets were competitive and the conditions for equation (3.1) held, our model (corollary II) would predict that Indonesia and China could not export to Baltimore at the same time as to LA. Thus, according to our theoretical model, market power must be present in LA market[38]. The statistical evidence seems to confirm that.

Baltimore-Los Angeles

[36] The average c.i.f. unit value of Peru in Los Angeles is also often higher than in New Orleans and the confidence interval for the median LOP deviation between New Orleans and LA would be negative for a 5% level of confidence.
[37] This is rather surprising, bearing in mind that ships have to cross the Panama Canal to go from Peru to New Orleans or to Baltimore.
[38] According to the analysis of equation (3.5), market power implies that one exporter drives rivals from the market through limit pricing. However, we cannot observe these instantaneous monopoly positions because of our monthly data.

We have also tested equation (4.1) for Baltimore (j) and Los Angeles (i), considering only the observations when China (m) was the arbitrageur. Thus, both independent variables refer to exports from China. The regression yields the following results:

lopdev = -0.031895 - 0.358485 tradecost - 0.077932 fobuvch

(0.007905)	(0.118713)	(0.132394)	(standard error)
[-4.034517]	[-3.019764]	[-0.588641]	[t-statistic]

We note that the intercept is significant, reinforcing our conclusion that exporters from China exercise market power in Los Angeles. The coefficient of the trade cost differential variable is negative, since the trade cost from China to Los Angeles is smaller than to Baltimore. Shipping time does not seem to help explaining lop deviations. Our attempts to correct for heteroscedasticity have not changed the above results in any significant way.

The behavior of the New York market is strongly marked by its decline over time. Bolivia ceased to export in September 1996, while Indonesia did so in September 2001. Indonesia was the largest exporter to New York in the 1989 to 2007 period (28% of total volume), but Brazil, the second largest (23%), was the most frequent exporter, present in 158 months compared with 152 months of Indonesia. Bolivia and China accounted for 19% and 15% of total imported volume at this destination, but China was present in 111 months, whereas Bolivia did so in 83 months. Trade costs to New York from the main exporters show a clear rise, starting in 1996. Generally, we cannot reject the hypothesis that LOP deviations between New York and other destinations are zero, according to the confidence intervals shown in Table (6). C.i.f. unit values from Indonesia are an exception. They tend to be slightly but persistently higher in New York as compared to Los Angeles. In fact, we cannot reject the hypothesis that Indonesia's and Brazil's c.i.f. unit values in New York are the same. Thus, bearing in mind that Indonesia reveals no market power in LA either, Indonesia's higher c.i.f. unit values in New York compared to LA result from Indonesia's lower trade costs to this latter destination.

6. Concluding Remarks

An international multi-country arbitrage model was developed in this paper to provide a theoretical basis for testing the law of one price and examining the relationship between LOP deviations and trade costs for a homogeneous and highly traded commodity. LOP deviations between the U.S. and Japan were shown to be as small as deviations between the two main exporters either in Japan or in the United States. LOP deviations between the U.S. and Japan have also revealed zero-mean reversion. Moreover, the response of unit values in one country to changes in unit values in the other country has shown to be quite fast (2 to 3 months), eliminating the possibility of persistently large deviations. In fact, the adjustment speeds we find between spatially separated markets are much faster than what has been found in previous studies, both intra-nationally and internationally.

These results provide strong support for the law of one price in the tin trade and confirm empirically the limited relevance of Heckscher's commodity points in understanding the link between trade costs and price differentials of location pairs. If we only consider the five largest exporters to the U.S. and the four largest to Japan, LOP deviations between the two importing countries and the speed of response to each other's price decline significantly. This suggests that fringe competitors in the U.S. and in Japan tend to raise deviations and slow down price convergence, making the arbitrage mechanism less efficient. Therefore, although competition (more competitors) tends to help to integrate markets, the presence of small competitors may delay the arbitrage mechanism between spatially separated markets.

LOP deviations between pairs of the main ports in the U.S. tend to be larger than between the U.S. and Japan, especially when the small markets of Los Angeles and San Francisco are one of the destinations. Therefore, there is no border effect in the tin trade between the U.S. and Japan, regardless of the behavior of exchange rates. This suggests that, without border-related trade barriers, local distributional costs and product differentiation, the border effect may vanish.

Among the main U.S. ports, LOP deviations are the smallest between Baltimore and New Orleans. We find strong evidence of frequent arbitrage by exporters from Peru and Brazil between these two destinations and no evidence of market power in these markets. In fact, the regularly large demand for imported tin and the presence of quite a few exporting countries shipping almost every month to Baltimore is an evidence of strong market competition. On the other hand, we find evidence of market power by exporters of China and by shipping services from Peru in Los Angeles. This explains some large LOP deviations between Los Angeles and other destinations in the U.S. Indeed, our theoretical model predicts that China would not sell to a lower price market (corollary II) such as Baltimore where trade cost is higher than LA, if the higher price market (LA) were competitive. The fact that it does must imply that the LA market is not competitive.

References

Anderson JE, van Wincoop E (2004) Trade Costs. Journal of Economic Literature XLII: 691-751.
Ardeni PG (1989) Does the law of one price really hold for commodity prices? American Journal of Agricultural Economics 7: 661-669.
Asplund M, Friberg R (2001) The Law of One Price in Scandinavian Duty-Free Stores The American Economic Review 91: 1072-1083.
Baffes J (1991) Some further evidence on the Law of One Price: the law of one price still holds. American Journal of Agricultural Economics 73: 1264-1273.
Ceglowski J (1994) The Law of One Price Revisited: New Evidence on the Behaviour of International Prices. Economic Inquiry 32: 407-418.
Crucini MJ, Shintani M (2008) Persistence in law of one price deviations: Evidence from micro-data. Journal of Monetary Economics 55: 629-644.
Cumby, RE (1996) Forecasting Exchange Rates and Relative Prices with the Hamburger Standard: Is what you want what you get with McParity? NBER Working Paper W5675.
Engel C, Rogers J (1996) How Wide is the Border? The American Economic Review 86: 112-1125.
____ (2001) Deviations from the law of one price: Causes and Welfare Costs, Journal of International Economics 55: 21-57.
Giovannini A (1988) Exchange Rates and Traded Goods Prices. Journal of International Economics 24: 45-68.
Goldberg PK, Knetter MM (1997) Goods Prices and Exchange Rates: what have we learned? Journal of Economic Literature 35: 1243-1272.
Goldberg PK, Verboven F (2001) The Evolution of Price Dispersion in the European Car Market. Review of Economic Studies 68: 811-848.
____ (2005) Market Integration and Convergence to the Law of One Price: Evidence from the European car market. Journal of International Economics, 65: 49-73.
Goodwin BK, Grennes T, Wohlgenant MK (1990) Testing the law of one price when trade takes time. Journal of International Money and Finance 9: 21-40.
Haskel J, Wolf H (2001) The Law of One Price - a case study. Scandinavian Journal of Economics 103: 545-558.
Hummels D, Lugovskyy V, Skiba, A (2008) The trade reducing effects of market power in international shipping. Journal of Development Economics: forthcoming.
Isard P (1977). How far can we push the law of one price? American Economic Review 67: 942-948.
Knetter MM (1989) Price Discrimination by U.S. and German exporters. The American Economic Review 79: 198-210.
____ (1993) International Comparisons of Pricing-to-Market Behaviour. The American Economic Review 83: 473-486.
Kravis IB, Lipsey RE (1978) Price behaviour in light of balance of payments theories. Journal of International Economics 8: 193-246.
Michael P, Nobay AR, Peel D (1994) Purchasing Power Parity Yet Again: Evidence from Spatially Separated Commodity Market. Journal of International Money and Finance 1: 637-657.
Obstfeld M, Taylor AM (1997) Nonlinear Aspects of Goods-Market Arbitrage and Adjustment: Heckscher's Commodity Points Revisited. Journal of the Japanese and International Economies 11: 441-479.
Parsley DC, Wei S (2001) Explaining the border effect: The role of exchange rate variability, shipping costs, and geography. Journal of International Economics 55: 87-105.
Protopapadakis AA, Stoll HR (1983) Spot and Futures Prices and the Law of One Price. Journal of Finance 38: 1431-1456.
____ (1986) The Law of One Price in International Commodity Markets: A reformulation and some formal tests. Journal of International Money and Finance 5: 335-360.
Richardson JD (1978) Some empirical evidence on commodity arbitrage and the law of one price. Journal of International Economics 8: 341-351.
Rogoff K (1996) The Purchasing Power Parity Puzzle. Journal of Economic Literature XXXIV: 647-668.
Thursby MC, Johnson PR, Grennes TJ (1986) The law of one price and the modelling of disaggregated trade flows. Economic Modelling 3: 293-302.
Vataja J (2000) Should the Law of One Price be Pushed Away? Evidence from International Commodity Markets. Open Economies Review 11: 399-415.
Wonnacot TH, Wonnacott RJ (1972) Introductory Statistics for Business and Economics. John Wiley & Sons, Wiley International Edition, New York

TABLE (1): Import unit value c.i.f. differentials of the two main exporters at each of the five largest destination ports in the U.S. (in log) from 1989 to 2007					
	Baltimore	New York	San Francisco	Los Angeles	New Orleans
median	0.0022	-0.0002	0.0031	-0.0088	-0.0033
mean	0.0020	-0.0061	0.0003	-0.0031	-0.0080
std dev	0.0584	0.0540	0.0719	0.0781	0.0621
max	0.1853	0.1981	0.1578	0.2137	0.1116
min	-0.1582	-0.1537	-0.1891	-0.1822	-0.1611
absdev mean	0.0440	0.0372	0.0544	0.0584	0.0505
observations	193	144	88	105	42
population	201	150	111	112	41
outliers	8	6	23	7	1

If the import unit values from the two main exporters are a and b, the import unit value differential at each month is calculated as log(a)-log(b) and the absolute deviation (absdev) is |log(a)-log(b)|

Table (2): C.i.f. import unit value differentials of the main exporters to Japan (in log) from 1989 to 2007							
	Leader	Indonesia	Indonesia	Indonesia	China	China	Thailand
	Follower	China	Thailand	Malaysia	Thailand	Malaysia	Malaysia
median	0.0010	-0.0066	-0.0106	-0.0105	-0.0058	-0.0061	0.0006
mean	0.0034	-0.0035	-0.0078	-0.0100	-0.0046	-0.0065	-0.0022
std dev	0.0305	0.0375	0.0281	0.0246	0.0426	0.0427	0.0303
max	0.1541	0.1541	0.1310	0.1247	0.1741	0.1656	0.1185
min	-0.1111	-0.1713	-0.0906	-0.0824	-0.1127	-0.1147	-0.1171
absdev mean	0.0222	0.0278	0.0205	0.0193	0.0320	0.0326	0.0208
observations	228	228	227	228	227	228	227

Table (3): C.i.f. import unit value differentials between the U.S. and Japan (in log) from 1989 to 2007							
Location Pairs	USA-Japan	US 5-Jap 4	US East-Jap 4	US West-Jap 4	US East-US West	BA-JAP	LA-JAP
median	-0.0020	-0.0054	-0.0065	0.0036	-0.0094	-0.0061	0.0065
mean	-0.0014	-0.0055	-0.0079	0.0064	-0.0142	-0.0035	0.0156
std dev	0.0349	0.0286	0.0324	0.0590	0.0625	0.0403	0.1050
max	0.0947	0.0882	0.0944	0.2123	0.2412	0.1848	0.6624
min	-0.1550	-0.1206	-0.1757	-0.2456	-0.3000	-0.1118	-0.4813
absdev mean	0.0259	0.0222	0.0242	0.0386	0.0410	0.0303	0.0592
observations	228	228	228	226	226	228	228

USA and Japan are series of c.i.f. unit values for the whole countries;
US 5 and Jap 4 include only the five and four main exporters to these countries, respectively;
US East includes the same 5 main exporters to the US, but only Baltimore and New York;
US West includes the same 5 main exporters to the US, but only to Los Angeles and San Francisco;
BA stands for Baltimore, includes all exporters and two missing values were replaced by the unit values of New York; and
LA stands for Los Angeles, includes all exporters and missing values were replaced by the unit values of San Francisco.

Table (4): LOP Deviations: unit values of the aggregate of all exporters to each destination					
	Baltimore	Baltimore	Baltimore	Baltimore	Los Angeles
	Los Angeles	New Orleans	New York	San Francisco	New Orleans
Median	-0.0098	-0.0060	0.0027	0.0103	0.0089
Mean	-0.0238	-0.0062	0.0047	0.0184	0.0236
Std Dev	0.1058	0.0523	0.0817	0.1030	0.1041
Max	0.5169	0.2773	0.5666	0.5506	0.6247
Min	-0.6383	-0.2277	-0.3198	-0.2903	-0.3236
Observations	214	190	205	208	179
10%	BA-LA<0	BAL-NO<0	BAL-NY=0	BAL-SF>0	LA-NO>0
5%	BA-LA<0	BAL-NO<0	BAL-NY=0	BAL-SF>0	LA-NO>0
1%	BA-LA<0	BAL-NO=0	BAL-NY=0	BAL-SF=0	LA-NO>0
	Los Angeles	Los Angeles	New Orleans	New Orleans	New York
	New York	San Francisco	New York	San Francisco	San Francisco
Median	0.0071	0.0118	0.0001	0.0078	0.0081
Mean	0.0217	0.0387	0.0060	0.0281	0.0055
Std Dev	0.1215	0.1435	0.0919	0.1123	0.1117
Max	0.6768	0.6573	0.6027	0.5236	0.4830
Min	-0.4627	-0.4607	-0.2708	-0.3231	-0.6540
Observations	198	197	171	174	190
10%	LA-NY>0	LA-SF>0	NO-NY=0	NO-SF>0	NY-SF>0
5%	LA-NY>0	LA-SF>0	NO-NY=0	NO-SF=0	NY-SF>0
1%	LA-NY=0	LA-SF>0	NO-NY=0	NO-SF=0	NY-SF=0

Table (5): LOP Deviations at f.o.b. values between US Main Ports: 1989/2007 (non-parametric sign test for large samples)											
Origins	Destinations	median	mean	std dev	min	max	obs	Confid. interval		SL	Ho
China	BA-LA	-0.051	-0.060	0.115	-0.318	0.636	93	-0.072	-0.029	1%	REJ
Indonesia	BA-LA	0.001	-0.015	0.122	-0.690	0.332	119	-0.003	0.007	10%	OK
Peru	BA-LA	-0.007	0.008	0.282	-0.688	2.293	76	-0.010	0.000	10%	OK
Brazil	BA-NO	-0.005	-0.003	0.043	-0.108	0.211	57	-0.011	0.003	10%	OK
Peru	BA-NO	0.000	-0.002	0.028	-0.111	0.133	117	-0.001	0.004	10%	OK
Peru	NO-LA	-0.007	-0.029	0.099	-0.686	0.088	71	-0.013	0.000	5%	OK
Peru	NO-SF	0.000	-0.034	0.215	-1.615	0.093	61	-0.003	0.005	10%	OK
Bolivia	NY-BA	-0.005	-0.005	0.035	-0.126	0.064	43	-0.012	0.000	10%	OK
Brazil	NY-BA	-0.001	-0.002	0.038	-0.129	0.142	151	-0.005	0.002	10%	OK
China	NY-BA	-0.002	-0.020	0.188	-0.911	0.476	61	-0.018	0.010	10%	OK
Indonesia	NY-BA	0.002	0.003	0.062	-0.348	0.291	117	-0.002	0.005	10%	OK
Bolivia	BA-SF	0.000	-0.005	0.083	-0.262	0.170	80	-0.008	0.007	10%	OK
Indonesia	BA-SF	0.008	0.065	0.191	-0.646	0.638	99	-0.001	0.019	1%	OK
China	BA-SF	-0.007	0.006	0.095	-0.258	0.219	54	-0.029	0.004	10%	OK
Peru	BA-SF	0.000	-0.008	0.057	-0.292	0.129	67	-0.003	0.003	10%	OK
Indonesia	SF-LA	0.002	-0.026	0.168	-1.012	0.677	81	-0.004	0.005	10%	OK
China	SF-LA	-0.090	-0.063	0.260	-0.295	1.181	31	-0.161	-0.035	1%	REJ
Peru	SF-LA	-0.001	-0.076	0.381	-2.302	0.274	40	-0.013	0.005	1%	OK
Brazil	NY-NO	-0.005	0.002	0.056	-0.150	0.228	58	-0.012	0.002	10%	OK
Bolivia	NY-SF	-0.009	-0.028	0.102	-0.282	0.177	36	-0.050	0.007	10%	OK
Indonesia	NY-SF	-0.001	-0.006	0.089	-0.671	0.192	80	-0.006	0.002	10%	OK
China	NY-SF	-0.008	0.003	0.115	-0.253	0.366	42	-0.022	0.001	10%	OK
Indonesia	NY-LA	0.003	-0.008	0.104	-0.411	0.481	148	-0.001	0.007	10%	OK
China	NY-LA	-0.047	-0.049	0.228	-0.488	0.585	33	-0.216	0.050	1%	OK

Table (6): LOP Deviations at c.i.f. values between US Main Ports: 1989/2007 (non-parametric sign test for large samples)											
Origins	Destinations	median	mean	std dev	min	max	obs	confidence interval		SL	Ho
China	BA-LA	-0.0493	-0.0573	0.1147	-0.3566	0.6342	93	-0.070	-0.028	1%	RE
Indonesia	BA-LA	0.0077	-0.0107	0.1209	-0.6895	0.3341	119	-0.002	0.016	1%	OK
Peru	BA-LA	-0.0122	-0.0384	0.0978	-0.6895	0.1307	75	-0.022	-0.003	1%	RE
Brazil	BA-NO	-0.0052	-0.0046	0.0453	-0.1192	0.2103	57	-0.011	0.003	10%	OK
Peru	BA-NO	-0.0016	-0.0040	0.0287	-0.1031	0.1304	117	-0.003	0.002	10%	OK
Peru	NO-LA	-0.0091	-0.0340	0.0999	-0.6856	0.0883	71	-0.024	0.000	1%	OK
Peru	NO-SF	-0.0046	-0.0176	0.0655	-0.2873	0.0950	60	-0.009	0.000	10%	OK
Bolivia	NY-BA	-0.0073	-0.0034	0.0336	-0.1167	0.0695	43	-0.012	0.001	10%	OK
Brazil	NY-BA	0.0004	-0.0010	0.0372	-0.1236	0.1371	151	-0.004	0.003	10%	OK
China	NY-BA	-0.0036	-0.0145	0.1871	-0.9268	0.4789	61	-0.014	0.012	10%	OK
Indonesia	NY-BA	0.0044	0.0067	0.0619	-0.3379	0.2946	117	-0.001	0.009	1%	OK
Bolivia	BA-SF	-0.0082	-0.0094	0.0801	-0.2551	0.1635	80	-0.015	0.003	10%	OK
Indonesia	BA-SF	0.0095	0.0646	0.1917	-0.6645	0.6332	99	0.000	0.017	5%	OK
China	BA-SF	-0.0150	0.0044	0.0912	-0.2482	0.2031	54	-0.029	0.007	10%	OK
Peru	BA-SF	-0.0057	-0.0186	0.0592	-0.2907	0.1249	67	-0.016	0.000	5%	OK
Indonesia	SF-LA	0.0077	-0.0221	0.1653	-0.9811	0.6869	81	-0.002	0.014	5%	OK
China	SF-LA	-0.0833	-0.0974	0.1219	-0.2887	0.2671	29	-0.151	-0.038	1%	RE
Peru	SF-LA	-0.0054	-0.0250	0.1262	-0.6928	0.2687	39	-0.025	0.000	1%	OK
Brazil	NY-NO	-0.0051	0.0012	0.0600	-0.1419	0.2211	58	-0.012	0.001	5%	OK
Bolivia	NY-SF	-0.0092	-0.0278	0.1023	-0.2822	0.1770	36	-0.045	0.011	10%	OK
Indonesia	NY-SF	-0.0008	-0.0065	0.0891	-0.6711	0.1917	80	-5E-05	0.005	10%	OK
China	NY-SF	-0.0076	0.0031	0.1153	-0.2525	0.3661	42	-0.017	0.005	10%	OK
Indonesia	NY-LA	0.0031	-0.0080	0.1045	-0.4113	0.4815	148	0.003	0.010	1%	RE
China	NY-LA	-0.0467	-0.0486	0.2279	-0.4882	0.5846	33	-0.130	0.025	1%	OK

Appendix I: Normality Tests and Zero-Mean Hypothesis Tests
Software: GRETL 1.7.6cvs

Two Main Exporters to	Test for null hypothesis	
	normal distribution p-value	population mean = 0 Two-tailed p-value
Baltimore	**0.14828**	**0.6431**
New York	0.00009	**-**
Los Angeles	**0.18469**	**0.6821**
San Francisco	**0.61076**	**0.9691**
New Orleans	**0.76608**	**0.4149**

Exporting Countries to Japan	Test for null hypothesis of normal distribution: p-value
Leader-Follower	0.00000
Indonesia-China	0.00000
Indonesia-Thailand	0.00000
Indonesia-Malaysia	0.00000
China-Thailand	0.00002
China-Malaysia	0.00025
Thailand-Malaysia	0.00000

Deviations of f.o.b. Unit Values: Test for null hypothesis of			
Exporting country	Destinations	normal distribution p-value	Mean=0 Two-tailed p-value
Bolívia	Baltimore-New York	0,00430	
Bolívia	Baltimore-San Francisco	0,00282	
Brazil	Baltimore-New York	0,00000	
Brazil	Baltimore-New Orleans	0,09758	**0,1395**
Brazil	New York-New Orleans	0,00000	
Peru	Baltimore-New Orleans	0,00000	
Peru	New Orleans-Los Angeles	**0,15817**	**0,1893**
Peru	San Francisco-Los Angeles	0,00147	
Peru	Baltimore-San Francisco	0,00000	
Peru	Baltimore-Los Angeles	0,00054	
Peru	New Orleans-San Francisco	0,00000	
China	Baltimore-New York	**0,66601**	**0,05911**
China	Baltimore-New Orleans	0,04458	
China	Baltimore-San Francisco	**0,19414**	**0,3109**
China	Baltimore-Los Angeles	**0,40434**	4,506e-008
Indonesia	Baltimore-New York	0,00000	
Indonesia	Baltimore-San Francisco	0,00000	
Indonesia	Baltimore-Los Angeles	0,00003	
Indonesia	New York-San Francisco	0,00000	
Indonesia	New York-Los Angeles	0,00000	
Indonesia	San Francisco-Los Angeles	0,00061	

Deviations of c.i.f. Unit Values: Test for null hypothesis of normal distribution	
Destinations	p-value
USA-Japan	0,00000
US 5-JAP 4	0,00111
US East-JAP 4	0,00000
US West-JAP 4	0,00000
US East-US West	0,00000

Appendix II: Nonparametric Sign Tests and Zero-Median Hypothesis Tests:

New York:
Non-parametric sign test
Ho: median=0
No. of observations: n=144
Confidence level: 90% (α=0.10): -0.006≤v≤0.003; **Ho cannot be rejected**.
Trials: 144; Successes: 71; p-value= 0.47; Ho is accepted.

Deviations of C.i.f. Unit Values of Exporting Countries to Japan				
Exporting Countries	**Confidence Interval**		**Significance Level**	**Ho: median equal zero**
Leader-Follower	-0.0024	0.0040	10%	Not Rejected
Indonesia-China	-0.0119	-0.0018	1%	Rejected
Indonesia-Thailand	-0.0142	-0.0063	1%	Rejected
Indonesia-Malaysia	-0.0141	-0.0076	1%	Rejected
China-Thailand	-0.0133	0.0012	1%	Not Rejected
China-Malaysia	-0.0144	0.0012	1%	Not Rejected
Thailand-Malaysia	-0.0022	0.0041	5%	Not Rejected

Deviations of C.i.f. Unit Values between U.S. Customs Districts				
Exporting Countries	**Confidence Interval**		**Significance Level**	**Ho: median equal zero**
Baltimore-Los Angeles	-0.0218	-0.0037	1%	Rejected
Baltimore-New Orleans	-0.0129	0.0004	1%	Not Rejected
Baltimore-New York	-0.0008	-0.0056	10%	Rejected
Baltimore-San Francisco	-0.0037	0.0258	1%	Not Rejected
Los Angeles-New Orleans	0.0006	0.0216	1%	Rejected
Los Angeles-New York	-0.0012	0.0157	1%	Not Rejected
Los Angeles-San Francisco	0.0034	0.0236	1%	Rejected
New Orleans-New York	-0.0057	0.0057	10%	Not Rejected
New Orleans-San Francisco	-0.0017	0.0194	5%	Not Rejected
New York-San Francisco	-0.0021	0.0169	1%	Not Rejected

Appendix III: ADF Tests

(E-Views 6.0)

Null Hypothesis: variable has a unit root

C.i.f. Unit Values or C.i.f. Deviations	Lag	Prob.*
Baltimore exporting leader	0	0.9563
Baltimore exporting follower	1	0.9832
Baltimore:leader-follower	1	0.0000
Baltimore: all exporters	0	0.9907
Los Angeles: all exporters	4	0.8340
New Orleans: all exporters	1	0.9895
New York: all exporters	1	0.9193
San Francisco: all exporters	1	0.9873
Baltimore-New York: all exporters	2	0.0000
Baltimore-New Orleans: all exporters	0	0.0003
Baltimore-Los Angeles: all exporters	0	0.0000
Baltimore-San Francisco: all exporters	1	0.0000
New York-New Orleans: all exporters	0	0.0000
New York-Los Angeles: all exporters	0	0.0000
New York- San Francisco: all exporters	0	0.0000
New Orleans-Los Angeles: all exporters	0	0.0000
New Orleans-San Francisco: all exporters	0	0.0000
Los Angeles-San Francisco: all exporters	0	0.0000
Japan exporting leader	2	0.8149
Japan exporting follower	0	0.9948
Japan: Leader-Follower	0	0.0000
Japan: China	0	0.9939
Japan: Indonesia	2	0.7986
Japan: Malaysia	1	0.9292
Japan: Thailand	1	0.9332
Japan: China-Indonesia	1	0.0000
Japan: China-Malaysia	1	0.0001
Japan: China-Thailand	1	0.0000
Japan: Indonesia-Malaysia	0	0.0000
Japan: Indonesia-Thailand	1	0.0000
Japan: Malaysia-Thailand	0	0.0000
USA	2	0.9512
Japan	1	0.9094
US 5	1	0.9459
JAP 4	1	0.9083
US East	1	0.9643
US West	0	0.8624
USA-Japan	1	0.0000
US 5-Jap 4	0	0.0000
US East-Jap 4	0	0.0000
US WEST_JAP 4	0	0.0000
US East-US West	0	0.0000
BA_Japan	0	0.0000
LA-Japan	0	0.0000

Exogenous: Constant

Lag length (Automatic based on SIC, MAXLAG=14)

*MacKinnon (1996) one-sided p-values.

Appendix IV: Cointegration Tests, VECM and Impulse and Response Tests
(E-views 6.0)
Endogenous variables are either c.i.f. unit values or c.i.f. unit value differentials
Trend assumption: No deterministic trend (restricted constant)
Series: Baltimore Leader BA1RST and Baltimore Follower BA2ND
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.000217; Probability 0.988240
Constant C= -0.005582 (0.00583) [-0.95765]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 0.870207; Probability 0.350898
Impulse Response Test (Cholesky Ordering: BA1RST BA2ND)
Response of BA1rst to BA2nd: Half-life=1 month; Response of BA2nd to BA1rst: Half-life=1 month
Exporting countries to Japan
Series: LEADER - FOLLOWER
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.360079; Probability 0.548462
Constant C= -0.003914 (0.00228) [-1.71592]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 2.502568; Probability 0.113661
Impulse Response Test (Cholesky Ordering: LEADER FOLLOWER)
Response LEADERtoFOLLOWER: Half-life=1 month; Response FOLLOWERtoLEADER: Half-life=1 month
Series: CHINA INDONESIA
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 1.741700; Probability 0.186924
Constant C= -0.002966 (0.00388) [-0.76432]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 0.890716; Probability 0.345283
Impulse Response Test (Cholesky Ordering: LEADER INDONESIA)
Response of CHINA to INDONESIA: Half-life=1 month; Response of INDONESIA to CHINA: Half-life=1 month
Series: CHINA MALAYSIA
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 3.509198; Probability 0.061029
Constant C= 0.007975 (0.00762) [1.04670]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 0.554307; Probability 0.456563
Impulse Response Test (Cholesky Ordering: CHINA MALAYSIA)
Response of CHINA to MALAYSIA: Half-life=1 month; Response of MALAYSIA to CHINA: Half-life=1 month
Series: CHINA THAILAND
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.684879; Probability 0.407912
Constant C= 0.007341 (0.00554) [1.32392]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 1.339258; Probability 0.247165
Impulse Response Test (Cholesky Ordering: CHINA THAILAND)
Response of CHINA to THAILAND: Half-life=1 month; Response of THAILAND to CHINA: Half-life=1 month
Series: INDONESIA MALAYSIA
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 8.809900; Probability 0.002996
Constant C= 0.010442 (0.00221) [4.72797]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 12.86144; Probability 0.000335
Impulse Response Test (Cholesky Ordering: INDONESIA MALAYSIA)
Response INDONESIA to MALAYSIA: Half-life=1 mth.; Response MALAYSIA to INDONESIA: Half-life=1 mth.
Series: INDONESIA THAILAND
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.293103; Probability 0.588239
Constant C= 0.009006 (0.00229) [3.94100]

VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 13.34379; Probability 0.000259
Impulse Response Test (Cholesky Ordering: INDONESIA THAILAND)
Response INDONESIA to THAILAND: Half-life=1 mth.; Response THAILAND to INDONESIA: Half-life=1 mth.
Series: MALAYSIA THAILAND
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 12.11477; Probability 0.000500
Constant C= -0.000747 (0.00275) [-0.27204]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 0.055094; Probability 0.814425
Impulse Response Test (Cholesky Ordering: MALAYSIA THAILAND)
Response MALAYSIA to THAILAND: Half-life=1 mth.; Response THAILAND to MALAYSIA: Half-life=1 mth.
UNITED STATES-JAPAN
Series: USA JAP
Lags interval (in first differences): 1 to 3
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 2.836608; Probability 0.092139
Constant C= -0.000111 (0.00308) [-0.03600]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 0.071862; Probability 0.036287
Impulse Response Test (Cholesky Ordering: USA JAP)
Response of USA to JAP: Half-life=3 months); Response of JAP to USA: Half-life=2 months
Series: US 5 JAP 4
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.437190; Probability 0.508482
Constant C= 0.005368 (0.00258) [2.08186]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 4.383589; Probability 0.788645
Impulse Response Test (Cholesky Ordering: US 5 JAP 4)
Response of US 5 to JAP 4: Half-life=2 months; Response of JAP 4 to US 5: Half-life=1 months
Series: US EAST JAP 4
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.454684; Probability 0.500119
Constant C= 0.007246 (0.00538) [2.36494]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 5.463439; Probability 0.019418
Impulse Response Test (Cholesky Ordering: US EAST JAP 4)
Response of US EAST to JAP 4: Half-life=2 months; Response of JAP 4 to US EAST: Half-life=1 month
Series: US WEST JAP 4
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.443600; Probability 0.505390
Constant C= -0.005469 (0.00306) [-1.01592]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 1.207541; Probability 0.271820
Impulse Response Test (Cholesky Ordering: US WEST JAP 4)
Response of US WEST to JAP 4: Half-life=1 month; Response of JAP 4 to US WEST: Half-life=1 month
Series: US EAST US WEST
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 levelVECM with linear restrictions (1,-1): Chi-square(1) 0.599809; Probability 0.438651
Constant C= 0.011953 (0.00595) [2.00725]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 4.086404; Probability 0.043230
Impulse Response Test (Cholesky Ordering: US EAST US WEST)
Response of US EAST to US WEST: Half-life=1 month; Response of US WEST to US EAST: Half-life=1 month
Baltimore-Japan: Series: BAL JAP
Lags interval (in first differences): 1 to 3
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 2.343938; Probability 0.125771
Constant C= 0.001499 (0.00371) [0.40350]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 0.037095; Probability 0.847271

Impulse Response Test (Cholesky Ordering: BAL JAP)
Response of BAL to JAP: Half-life=4 months; Response of JAP to BAL: Half-life=1 month
Los Angeles-Japan: Series: LA JAP
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Chi-square(1) 0.541494; Probability 0.461814
Constant C= -0.015243 (0.00726) [-2.09982]
VECM with linear restrictions (1,-1) without intercept: Chi-square(1) 3.673999; Probability 0.055267
Impulse Response Test (Cholesky Ordering: LA JAP)
Response of LA to JAP: Half-life=1 month; Response of JAP to LA: Half-life=2 months
US PORTS (ALL EXPORTERS)
Baltimore – New Orleans: Series: BA NO
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.927644; Constant C= 0.004205 (0.00518) [0.81230]
VECM with linear restrictions (1,-1) without intercept: Probability 0.425333
Impulse Response Test (Cholesky Ordering: BA NO)
Response of BA to NO: Half-life=1 month; Response of NO to BA: Half-life=1 month
Baltimore- New York: Series: BA NY
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.686203; Constant C= -0.003819 (0.00612) [-0.62443]
VECM with linear restrictions (1,-1) without intercept: Probability 0.498136
Impulse Response Test (Cholesky Ordering: BA NY)
Response of BA to NY: Half-life=1 month); Response of NY to BA: Half-life=1 month
Baltimore – San Francisco: Series: BA SF
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.997845; Constant C= -0.014566 (0.00997) [-1.46029]
VECM with linear restrictions (1,-1) without intercept: Probability 0.164817
Impulse Response Test (Cholesky Ordering: BA SF)
Response of BA to SF: Half-life=1 month; Response of SF to BA: Half-life=1 month
New York – New Orleans: Series: NY NO
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.652495; Constant C= 0.003951 (0.00802) [0.49255]
VECM with linear restrictions (1,-1) without intercept: Probability 0.577419
Impulse Response Test (Cholesky Ordering: NY NO)
Response of NY to NO: Half-life=1 month; Response of NO to NY: Half-life=1 month
New York – San Francisco: Series: NY SF
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.042255; Constant C= 0.003780 (0.01201) [0.31466]
VECM with linear restrictions (1,-1) without intercept: Probability 0.566225
Impulse Response Test (Cholesky Ordering: NY SF)
Response of NY to SF: Half-life=2 months; Response of SF to NY: Half-life=1 month
New Orleans – San Francisco: Series: NO SF
Lags interval (in first differences): 1 to 1
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.902228; Constant C= -0.022716 (0.01820) [-1.24800]
VECM with linear restrictions (1,-1) without intercept: Probability 0.240559
Impulse Response Test (Cholesky Ordering: NO SF)
Response of NO to SF: Half-life=1 month; Response of SF to NO: Half-life=1 month
Baltimore – Los Angeles: Series: BA LA
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.218795; Constant C= 0.020965 (0.00764) [2.74239]
Impulse Response Test (Cholesky Ordering: BA LA)

Response of BA to LA: Half-life=1 month; Response of LA to BA: Half-life=1 month
Los Angeles - New Orleans: Series: LA NO
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.391369; Constant C= -0.024032 (0.00854) [-2.81489]
Impulse Response Test (Cholesky Ordering: LA NO)
Response of LA to NO: Half-life=1 month; Response of NO to LA: Half-life=1 month
Los Angeles – San Francisco: Series: LA SF
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.222872; Constant C= -0.033894 (0.01360) [-2.49180]
Impulse Response Test (Cholesky Ordering: LA SF)
Response of LA to SF: Half-life=1 month; Response of SF to LA: Half-life=1 month
New York-Los Angeles: Series: NY LA
Lags interval (in first differences): No lags
Unrestricted Cointegration Rank Test (Trace): Trace test indicates 1 cointegrating eqn(s) at the 0.01 level
VECM with linear restrictions (1,-1): Probability 0.849332; Constant C= 0.021302 (0.00863) [2.46742]
VECM with linear restrictions (1,-1) without intercept: Probability 0.425333
Impulse Response Test (Cholesky Ordering: NY LA)
Response of NY to LA: Half-life=1 month; Response of LA to NY: Half-life=1 month
Summary (U.S. ports)
VECM: Test of Restriction (1,-1)

Chi-square(1)	Probability	Constant	t-statistics	Implication	Half-life
0.008246	0.927644	0.004205	[0.81230]		
0.635534	0.425333	without constant		BA≈NO	1 month
0.163227	0.686203	-0.003819	[-0.62443]		
0.458905	0.498136	without constant		BA≈NY	1 month
7.29E-06	0.997845	-0.014566	[-1.46029]		
1.929462	0.164817	without constant		BA≈SF	1 month
0.202769	0.652495	0.003951	[0.49255]		
0.310426	0.577419	without constant		NY≈NO	1 month
4.124973	0.042255	0.003780	[0.31466]		
0.329039	0.566225	without constant		NY≈SF	1 month
0.015092	0.902228	-0.022716	[-1.24800]		
1.377313	0.240559	without constant		NO≈SF	1 month
1.512256	0.218795	0.020965	[2.74239]	BA<LA	1 month
0.734683	0.391369	-0.024032	[-2.81489]	NO<LA	1 month
1.485784	0.222872	-0.033894	[-2.49180]	SF<LA	1 month
0.036089	0.849332	0.021302	[2.46742]	NY<LA	1 month