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State Space Application to Recent Automobile Sector Triangle Trade between Japan and Latin America*

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Abstract

The flow of foreign direct investment (FDI) has the effect of shifting trade patterns, especially those of developing countries. With FDI flow growth, vertical intra-industry trade increases as those production-fragmentation-led investments mature. This relation is readily apparent in the respective automobile sectors of Brazil and Mexico. This paper generates empirical evidence on the vertical intra-industry trade pattern between Japan-Brazil and Japan-Mexico using a state space econometric framework. Results show that Mexico is more integrated into the worldwide production chain than Brazil, where risks such as exchange rate volatility are higher. Results also show how each structure is affected by the 2008 recession and how the difference in the intensity can be regarded as further evidence of how Mexico is advancing more rapidly into the fragmentation process than Brazil is.

Keywords: Triangle trade, State space models, Exchange rate volatility, Automobile industry

JEL Classification: F14

1 Introduction

Seminal work by Jones et al. (2005) demonstrated that production fragmentation occurs when a sufficiently large factor price differential prevails between countries. Thereby, firms are able to reduce marginal costs by splitting production units that use production factors in distinct intensities. Fragmentation also demands that service link costs be sufficiently low for services such as transportation and communication, enabling them to organize multi-plant operations concurrently. These costs are of a fixed kind. Therefore, scale economies are crucially important.

Low wages, inexpensive infrastructure services, industrial agglomeration, and a favorable policy environment are set as basic characteristics that firms are seeking when setting up production in another country. This is done through foreign direct investment (FDI). As the fragmentation-related investments made by foreign firms begin to consolidate, trade between the FDI origin country and the FDI receptor also increase as a sign of integration into the worldwide production structure.

Asian countries such as China, Thailand, and Singapore have such characteristics. For them, those characteristics can explain part of their good FDI performance, according to the UNCTAD performance index. Japanese firms play a crucial role as major investors in those economies, providing capital and technology to boost production and organizing inter-firm relationships (Kimura and Ando, 2003). Consequently, the links between the Japan economy and those of Asian countries are complex, linked by many sector FDI flows and intra-industry trade (Kimura et al., 2007).

Empirical studies related to this subject have mainly investigated Asia, especially the electronics industries in Asian countries. Yi (2003), Fukao et al. (2003), Xing (2007), Thorbecke (2008a), Dean et al. (2009), Sawyer et al. (2010), and Ferrarini (2013) all present empirical evidence demonstrating how interconnected and important Japanese exports to China and Chinese exports to the United States are. They also reveal that Japan is increasing its exports to the United States in an indirect manner. This fact might not be readily apparent because exports are measured by gross value rather than on a value-added basis, leading to an underestimation of intermediate goods exports.

In contrast, Latin America has not experienced development of its production and distribution networks to the same degree that Asian economies have. Because of import-substitution led economic policy agendas that have dominated economic policy in the region for a long time, Latin America's insertion into international markets has been more limited, therefore limiting the opportunities inherent in globalization. In the early 1990s, however, a major shift occurred in economic policy orientation aimed at opening and integrating Latin America economies into a globalizing world. With a higher degree of openness, automobile firms, expecting to improve access to South American markets from Brazil and to North American market from Mexico, began to invest more in the region (Mortimore,2000; Treviño et al., 2004). Diehl (2001), Spatz and Nunnenkamp (2004), and Faustino and Leitão (2009) present empirical evidence that fragmentation became a relevant subject also in the automobile sector with increased participation of Latin America economies, especially of Mexico and Brazil (Lall et al., 2010).

The automobile sector exhibits extremely strong links between Japan and Latin America that date back to the 1950s. According to data from JETRO, Japanese FDI stock in Brazil and Mexico in 2011 represented nearly 20% of the total in the region. The automobile sector is extremely important because the numbers of firms in the segment increased rapidly after 2003 (Toyo Keizai Shinposha 2012). In 2012, Japanese automobile company production was about 800 thousand cars in Mexico and 450 thousand in Brazil, which seems rather modest compared to the more than 2 million produced in China and Thailand. However, Japanese firms have announced aggressive plans for expansion in Mexico and Brazil. The two countries are expected to grow into regional production hubs. Parts and components for automobile production have become predominantly important export items from Japan to each country. Despite their growing importance, Japanese automobile companies' involvement in Latin American economies has rarely been studied using a rigorous statistical method.

The main objective of this study is to evaluate triangular trade from Japan to Mexico or Brazil, and then to the remainder of the world. Our main emphasis is to ascertain whether observed conditions are amenable to production fragmentation or not. We apply a state space econometric framework, which can accommodate structural breaks and time-varying coefficients, characteristics which are desirable because Latin American economies have endured economic-policy-driven structural shocks.

The following sections present a brief summary of the recent literature, a description of the econometric framework, estimation results, and salient conclusions.

2 Triangular Trade Pattern Empirical Model

As Jones et al. (2005) point out, service link costs constitute a key index to evaluate conditions for production fragmentation. Our approach is similar to that used by Thorbecke (2008a, 2008b) and Kimura and Ando (2003), who analyze the effects of exchange rate volatility on triangular trade patterns. Following these works, we assume that service link costs are an increasing function of risk and uncertainty caused by exchange rate volatility.

The relationship between service link costs and exchange rate volatility can be explained by the fact that exchange rate volatility impacts the importers' and exporters' decisions. Both the exporter and the importer are on opposite sides of a risky trading relationship. Their respective roles are reversed, so greater uncertainty related to exchange rates might affect their decisions related to their propensity to seek exposure to risky transactions, resulting in a lower level of international trade (Oskooee and Hegerty, 2007).

According to Dixit (1989), the exchange rate volatility has additional effects on trade because it influences companies' investment decisions and the fragmentation process itself. Because foreign firms must overcome cost barriers to enter a new market, greater exchange rate volatility makes the entry decision more costly: firms might be averse to abandoning their sunk investment when exchange rates change adversely. Because higher volatility can be expected to engender higher costs, companies tend to be more risk-averse with respect to countries that have high exchange rate volatility.

A large body of empirical literature describes exchange rate volatility's impact on international trade. Regarding the Asian evidence, Nishimura and Hirayama (2013) and Hayakawa and Kimura (2009) have produced two recent articles. Although their econometric frameworks mutually differ, both present evidence of a negative impact on trade by higher nominal exchange rate volatility. They emphasize aggregate trade volume and do not examine triangular trade patterns. Regarding Latin America, a cointegration analysis by Arize et al. (2008) provides evidence of a negative impact on aggregate trade by higher exchange rate volatility.

3 The Model and Data Issues

A straightforward means of evaluating the impact of exchange rate volatility on triangular trade was described by Thorbecke (2008a, 2008b), whose work we used as a reference to propose the following.

$$\begin{aligned} &\ln Cex_{jpn,it} = \gamma_t + \beta^1 \ln Fex_{i,Wt-1} + \beta^2 \ln Fex_{i,Wt-12} \\ &+\theta^1 \ln Rer_{t-1} + \theta^2 \ln Rer_{t-12} + \lambda^1 \hat{\sigma}_{t-1} + \lambda^2 \hat{\sigma}_{t-12} + \varepsilon_t \end{aligned} \tag{1}$$

In that equation, the following variables and terms were used.

$\ln Cex_{jpn,it}$ Natural logarithm of Japanese automobile component exports to country i at t

$\ln Fex_{i,Wt-1}$ Natural logarithm of country i final automobile exports to world at $t-1$

$\ln Fex_{i,Wt-12}$ Natural logarithm of country i final automobile exports to world at $t-12$

$\ln Rer_{t-1}$ Natural logarithm of country i real exchange rate at $t-1$ (country currency against the Japanese yen)

$\ln Rer_{t-12}$ Natural logarithm of country i real exchange rate at $t-12$

$\hat{\sigma}_{t-1}$ Estimated nominal exchange rate volatility for country i currency against the Japanese yen at $t-1$

$\hat{\sigma}_{t-12}$ Estimated nominal exchange rate volatility for country i currency against the Japanese yen at $t-12$

Before proceeding to the estimation, some caveats related to data are in order. First, these analyses use time series analysis rather than panel data analysis because of the limited size of the sample consisting of Brazil and Mexico as Japan's trade partner. The monthly frequency dataset is used because of its availability and short span of the dataset.

Second, we define final goods and components in the automobile sector as described in Table 1. The goods coverage is limited to those under Chapter 87 of the Harmonized System, although automobile production procures parts and materials from a much wider range of products.

Table 1: Description of Goods Used for Estimation

Code	Final goods
8702	Vehicles for people transportation
8703	Cars and other motor vehicles
8704	Vehicles for transport of goods
8705	Special purpose motor veichles
8711	Motorcycles
	Components
8706	Chassis with engines
8707	Bodies for vehicles
8708	Parts and accessories
8714	Accessories of vehicles

The source for Japanese exports to Brazil and Mexico is the Ministry of Finance of Japan[1]. The Brazilian final goods exports to world data source is the Ministry of Development[2]. The Mexican final automobile exports to the world data source is the Ministry of Finance[3]. The exchange rates and CPI data used to calculate the real exchange rate were obtained from the respective countries' central banks.

Third, sample sizes vary among countries. Mexican monthly trade data are available from January 2003, but they constitute a smaller sample than the data of Brazil, which are available from January 1997. For that reason, estimation is done for respective countries using different sample sizes.

The data availability is not the only reason to estimate models having different sample sizes, but to account for important economic changes that occurred in the region during the late 90s and early 2000s. For the Brazilian case, the period included important structural changes such as the end of the fixed exchange rate policy in 1999 and a more socially focused economic policy agenda with the Worker Party presidential administration coming into office in 2003.

During that same period in Mexico, the Mexican Peso crisis in 1994, NAFTA creation in 1995, and the bilateral trade agreement with Japan concluded in 2005 can be expected to have caused important structural breaks that are difficult to address because no data exist for periods before those events. At least those important changes can be analyzed in the Brazil case.

However, accommodating structural changes is not merely a question of data availability; it is also econometrically challenging. As an estimation framework, the space state model allows greater flexibility to address structural changes. The next section presents a review of the state space framework.

[1] `http://www.customs.go.jp/toukei/info/tsdl_e.htm`
[2] `http://aliceweb2.mdic.gov.br/`
[3] `http://www.economia.gob.mx/communidad-negocios/comercio-exterior/informacion-estadistica-y-arancelaria`

4 State Space Models

Although the application of state space models into econometrics can be traced back to the 1970s, with papers by Rosenberg (1973), Garbade (1977), and Harvey and Phillips (1979), the methodology is far from being applied to its full potential. Since the appealing features of space state models were first advocated in a report by Harvey (1997), state space models' utilization and related discussions have increased considerably.

The most positive aspect of the state space models is their greater flexibility compared to traditional models used for time series econometrics. The traditional approach relies heavily on identification procedures represented by unit root tests, cointegration tests, and the sample correlogram analysis. Under circumstances in which no sample is available to identify key data components such as cointegration rank or autocorrelation order, the estimated model might lack robustness. Moreover, a wide range of debate surrounds the lack of power in unit root tests, which represents an important aspect of the traditional analysis, especially in the presence of structural breaks.

State space models do not require the same identification procedure that a traditional model does. Unit root tests, which generate evidence related to stationarity of a given time series, are completely unnecessary because stationarity is not a condition for state space models. Consequently, state space models are more flexible with weaker conditions than traditional ones. However, for state space model estimation, the computational power demand is much higher than that for traditional models.

A salient feature of space state models is their capability of writing a given time series set as a function of unobserved components that can be estimated using a combination of the Kalman Filter and maximum likelihood estimator. Writing a given system in state space terms is necessary to specify how the unobserved components interact with the time series that is to be analyzed. Such writing is done using the observation equation, but it is also important to specify how those unobserved components evolve over time. In fact, this is accomplished using the state equation.

With the state and observational equation defined, state space estimation is done using a Kalman filter. The filter objective is to update knowledge of the system each time a new observation for the variable for which the analysis has been conducted becomes available. Filtering yields the Kalman filter residual and its variance, which is a one-step-ahead forecast error and a one-step-ahead forecast error variance given the sample information to $t-1$. The Kalman filter equations for the prediction error and its variance are also designated as the Kalman gain[4].

Although knowledge of the system dynamics is given by the Kalman filter, the estimation procedure is not described completely. In the filtering update equations, the state vector error variances are important components. Because all are unknown, they must be estimated.

To estimate the error variances, which is embodied by hyper-parameters in state space terminology,

[4] On this subject, Harvey (1989), Hamilton (1994), Kim and Nelson (1999), and Durbin and Koopman (2001) are excellent references.

the maximum likelihood estimator is applied in the prediction error ω_t and the prediction error variance $\mathbf{F}_t$ by the equation designated as the prediction error decomposition.

$$\ln L(y) = -\frac{n}{2} \ln 2\pi - \frac{1}{2} \sum_{t=1}^{n} \ln(\det \mathbf{F}_t) + \omega_t' \mathbf{F}_t^{-1} \omega_t \tag{2}$$

The quantities ω_t and $\mathbf{F}_t$ are calculated using the Kalman filter. The maximum likelihood is computed from the filter output. To initiate the filter, when $t = 1$, a question arises related to the initial conditions. The most common case is to set the state vector as 0 and to set its variance as a large number[5]. Because the Kalman filter is initiated and the sample observations bring it knowledge regarding the stochastic process that describes the data, those values converge to stable figures.

The space state methodology will be applied to estimate the triangular trade pattern but also to estimate the nominal exchange rate volatility. The next section describes how it was estimated.

5 Stochastic volatility in the space state framework

The exchange rate volatility is a major component of the triangular trade pattern. Although some discussion arises with respect to how the volatility must be measured to account for it as a regressor in the triangular trade pattern.

The most straightforward method of incorporating consideration of the volatility effect is applying the historical volatility as a regressand, but as Engle (1983) notes, the conditional variance is of greater relevance to economic agents planning their behavior. Considering that the nominal exchange rate is a difficult variable to forecast (Obstfeld and Rogoff, 2001), the expectation component in volatility might include a considerable amount of information related to economic agent behavior. That component justifies conditional volatility as a regressand.

The ARCH-GARCH-based models represent a logical and direct means to estimate the expected volatility. However, as is true for most time domain models, it relies on identification of the correlogram. If the sample correlogram analysis is insufficient to identify any ARCH-GARCH process, then its application might be severely compromised.

Harvey et al. (1994) and Durbin and Koopman (2001) propose a framework to estimate the volatility in a series using a space state model that is applicable in cases with no sample evidence that can justify an ARCH-GARCH model approach or where the estimated ARCH-GARCH models do not satisfy the fourth moment condition (Harvey et al., 1994). Although this framework is flexible, it becomes a Quasi Maximum Likelihood Estimator when dealing with volatility.

[5]The Oxmetrics 5 State space package, called Ssfpack 3.1, sets the initial value for the variance as high as 100,000, but such a value can be changed by any desired amount.

Considering x as the demeaned percentage variation of the nominal exchange rate, then the behavior can be modeled as

$$x_t = \sigma \varepsilon_t \exp\left(h_t/2\right) \tag{3}$$

Where

$$h_t = \varphi h_{t-1} + \eta_t \tag{4}$$

A linear approximation of equation (3) is shown below.

$$\ln x_t^2 = h_t + \kappa_t + \xi_t \tag{5}$$

The Kalman filter result for h_t is the estimated volatility for the given x series. Equation (5) represents the observation equation, whereas equation (4) is the transition equation and φ is the volatility persistence parameter.

For the estimation procedure, exchange rate data from January 1997 through December 2012 are available for both countries. Table 2 presents the estimated parameters for the stochastic volatility described in equations (3)-(5) for the Brazilian real and Mexican peso against the Japanese yen. Figure 1 shows the estimated volatility against time.

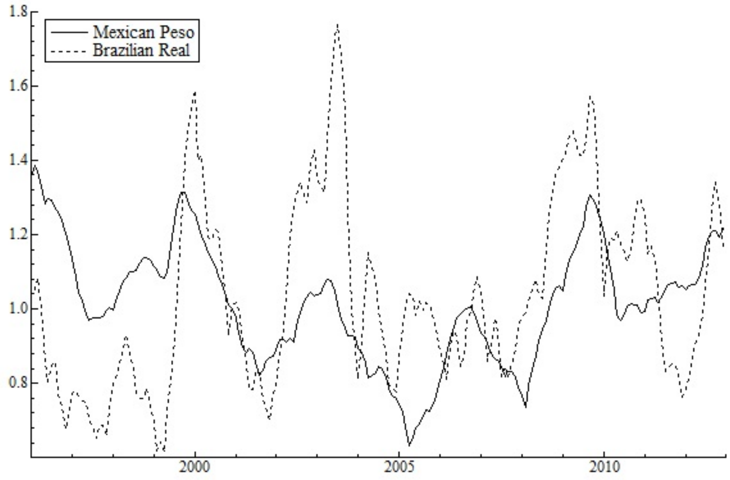
To facilitate interpretation and further readings, we chose to use the standard deviation instead of the variance. It can be read better with the same unity, percentage variation, than the average.

Table 2: Stochastic Volatility Estimated Parameters

Parameters	**Brazilian real**	**Mexican peso**
$\hat{\varphi}_i$	0.78	0.86
s.d.	0.32	0.2
$\hat{\sigma^2}_\eta$	0.69	0.39
LogL	-177.2	-271.27

According to the estimation results presented in Table 2, the exchange rate volatility between the Brazilian real and yen is higher than the Mexican peso - yen rate. This result derives from the higher error variance (0.69 for Brazil, 0.39 for Mexico case) and a lower φ, estimative for Brazil equation (0.78 for Brazil, 0.86 for Mexico). The lower the coefficient φ is in absolute terms, the lower is the temporal dependence effect which results in a more erratic pattern for volatility.

Figure 1: Estimated Nominal Exchange Rate Volatility



The vertical axis in Figure 1 represents the estimated standard deviation for the Brazilian and Mexican nominal exchange rates from the state space model presented in equations (3)-(5). The standard deviation is given in percentage terms, which means that the exchange rate has an average of X% deviation from the mean. Consequently, a higher number on the vertical axis can be read as an expectation that an exchange rate will be more volatile in the period shown on the horizontal axis.

During the sample period, volatility for both countries showed similar behavior in many respects. For example, volatility spikes tended to occur almost simultaneously for both currencies, but the shocks generate a more sensible response in Brazilian currency, especially after 1999, when Brazil ended the Real Plan and its fixed exchange rate policy.

After the end of the Real Plan, Brazilian exchange rate volatility increased systematically, which can be regarded as evidence that investors infer that Brazil as a riskier country than Mexico. That inference affects not only investment decisions, but also how Japanese companies behave with respect to each country.

The estimated exchange rate volatility is applied hereinafter in the triangular trade equations as an explanatory variable.

6 Estimations for the Triangular Trade Equations

The triangular trade pattern described in equation (1) can be written in state space specifications as the following.

$$y_t = \mathbf{Z}\Gamma_t + \varepsilon_t \tag{6}$$

$$\Gamma_{t+1} = \mathbf{T}\Gamma_t + R_t\mathbf{v}_t \tag{7}$$

$$
\begin{aligned}
&\gamma_{t+1} = \gamma_t + \Theta B + v_t^1 \\
&\beta_{t+1}^1 = \beta_t^1 + v_t^2 \qquad && \beta_{t+1}^2 = \beta_t^2 + v_t^3 \\
&\theta_{t+1}^1 = \theta_t^1 + v_t^4 \qquad && \theta_{t+1}^2 = \theta_t^2 + v_t^5 \\
&\lambda_{t+1}^1 = \lambda_t^1 + v_t^6 \qquad && \lambda_{t+1}^2 = \lambda_t^1 + v_t^7
\end{aligned} \tag{8}
$$

$$
\begin{aligned}
&y_t = \ln Cex_{jpn,it} \\
&\mathbf{Z} = \begin{bmatrix} 1 & \ln Fex_{i,Wt-1} & \ln Fex_{i,Wt-12} & \ln Rer_{t-1} & \ln Rer_{t-12} & \hat{\sigma}_{t-1} & \hat{\sigma}_{t-12} \end{bmatrix}
\end{aligned} \tag{9}
$$

$$
\Gamma_{t+1} = \begin{bmatrix} \gamma_{t+1} \\ \beta_{t+1}^1 \\ \beta_{t+1}^2 \\ \theta_{t+1}^1 \\ \theta_{t+1}^2 \\ \lambda_{t+1}^1 \\ \lambda_{t+1}^2 \end{bmatrix} \quad \mathbf{v}_t = \begin{bmatrix} v_t^1 \\ v_t^2 \\ v_t^3 \\ v_t^4 \\ v_t^5 \\ v_t^6 \\ v_t^7 \end{bmatrix} \quad Var(\mathbf{v}_t) = \begin{bmatrix} \sigma_{v^1}^2 & 0 & \dots & 0 \\ 0 & \sigma_{v^2}^2 & \dots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \dots & \sigma_{v^7}^2 \end{bmatrix} \quad \mathbf{T} = \mathbf{I}_7 \tag{10}
$$

Equation (1) represents the observation equation. Equations in eq. (8) are the state equations. The parameters to be estimated in the model are the β, θ and λ coefficients, and the variance for ε and all v error terms. The ΘB term represents n endogenous structural breaks in the stochastic trend, as described by Harvey and Koopman (1992).

The regression coefficients have an independent dynamic through time. If the variances for its errors are set as equal zero. Then the coefficients are constant over time. For each explanatory variable, two lags were added, the one month lag $(t-1)$ to avoid contemporaneous effects which results in simultaneous bias and the one year lag $(t-12)$, which is intended to capture dynamic effects because variables such as the real exchange rate might take more time to affect international trade volumes.

With the equation written in state space form, it is necessary to verify whether such a specification fits the data well. According to Harvey (1989), it is possible to determine a more appropriated specification using Likelihood Ratio tests (Harvey, 1989) or by Akaike Information Criteria (AIC; Harvey, 1989; Durbin and Koopman, 2001). Specification tests in state space models are done in the first step of the estimation procedure because the results will determine other model aspects such as structural breaks.

The first attempt is to estimate the model as specified in equations (1) and (7)-(10). It is described in such way that the trend is stochastic with possible structural breaks. There are six time variable coefficients and the observation equation error variance, resulting in eight hyperparameters (error variances) to be estimated. To show the most suitable specifications, the Likelihood test and AIC are displayed in Tables 3 and 4.

Although the AIC might be used as the only criterion to choose the specification, the AIC was not used alone because it tends to choose overparameterized models. Moreover, the result might generate state components that are not significantly different from zero, such as a regression coefficient that is set to vary over time. However, it might be no different from zero when the standard deviation is considered. If that is true, then setting the coefficient as fixed to reduce the hyperparameter vector is indicated (Harvey, 1989). The Likelihood Ratio test was applied to support the chosen specification.

For the Brazil equation, the specification choice was more direct because the AIC and the Likelihood Ratio tests point to the model with fewer hyperparameters to be estimated. The Mexican equation, as the first analysis suggests, has two models: the #2 model, according to AIC and #4. The latter was chosen as the LR test between model #2 and #4, with a *P-value* equal 1, suggesting that the model #4 describes the data as well as model #2, but with fewer hyperparameters.

Table 3: Specifications for Brazilian Equation

Model Number	**Model Description**	**Estimated Hyperparameters**	**LogL**	**SSE**	**AIC**	*P value* **LR test Against #1**	n
1	*Stochastic Trend* *2 variables coeff for:* *RER* *Final Goods exports* *Volatility*	8	293.1	0.0188	0.022		179
2	*Stochastic Trend* *T-1 lag* *effects are variable for:* *RER* *Final Goods exports* *Volatility*	5	263.23	0.027	0.029	1	
3	*Stochastic Trend* *T-12 lag effect* *are variable for:* *RER* *Final Goods exports* *Volatility*	5	283.95	0.023	0.0254	1	
4	**Chosen Model** *Stochastic Trend* *T-12 lag effect variable only on* *RER and Final Goods exports* *Fixed Effect for* *Exchange Rate Volatility*	4	270.84	0.024	0.0253	1	

Table 4: Specifications for Mexican Equation

Model Number	**Model Description**	**Estimated Hyperparameters**	**LogL**	**SSE**	**AIC**	*P value* **LR test Against #1**	n
1	*Stochastic Trend* *2 variables coeff for:* *RER* *Final Goods exports* *Volatility*	8	159.72	0.02	0.0264		
2	*Stochastic Trend* *T-1 lag* *effects are variable for:* *RER* *Final Goods exports* *Volatility*	5	159.41	0.023	0.027	1	
3	*Stochastic Trend* *T-12 lag effect* *are variable for:* *RER* *Final Goods exports* *Volatility*	5	159.28	0.024	0.0283	1	108
4	**Chosen Model** *Stochastic Trend* *T-12 lag effect variable only on* *RER and Fixed effect for* *Final Goods exports and* *Exchange Rate Volatility*	3	159.09	0.025	0.0274	1	

Tables 5-7 and the graphics in Figures 2-4 show results from the chosen specifications. Table 5 presents estimated hyperparameters (standard deviation for each state equation). Figures 2-4 portray the time-variant coefficients. Table 6 shows estimation results for fixed coefficients. Table 7 shows the structural break results.

Table 5: Estimated Hyperparameters

Hyperparameter	**Brazil**	**Mexico**
Observation error s.d.	0.0854	0.1285
Stochastic Trend error s.d.	0.0045	0.0014
T-12 lag RER coefficient equation s.d.	0.0173	0.0224
T-12 lag Final good exports coefficient equation error s.d.	0.0158	Not Estimated

The vertical axis of Figure 2 shows the estimated coefficient for world exports, β_t^2,and its 95% confidence interval (dotted line) for the Brazil equation. The result suggests that the magnitude by which components imports reacts from a variation in exports amount is changing over time. It can be regarded as evidence of how Brazil-based exporters are changing their decision-making processes. Because the specification in equation (1) describes both variables in natural logarithm,β_t^2 can be interpreted as an elasticity coefficient. Given a 1% increase in final goods exports, Brazilian components imports tend to rise β_t^2%.

A positive relation between world exports and component imports from Japan was expected because, as long as the exports are increasing, with no component substitution in the domestic market, imports will increase. Although exports are not the emphasis of Brazil-based producers, exports have been increasing, which can be regarded as evidence of Brazil's integration into the worldwide production structure.

Figure 2: World Export Coefficient for Brazilian Equation.

The idea that exporters seek to maintain stable export flows to keep costs low, combined with the Japanese "Just in Time" production paradigm and the fact that production in Mexico is exclusively intended for export can explain why final exports have no statistically significant impact in component imports for the Mexico equation.

This interpretation can be reinforced by results related to the real exchange rate coefficients. The correlation between real exchange rate devaluation and a boost in component imports was expected because it increases exchange rate led export growth. However this effect can take some time to become apparent because the only coefficient that is significantly different from zero is the 12-month lagged real exchange rate, θ_t^2.

The standard deviation for the real exchange rate coefficient equation is higher in Mexico (0.022 against 0.017 in Brazil) because the exchange rate devaluation impact is not constant over time, which engenders the conclusion that Mexico-based exporters are more eager to change mark-ups than Brazil-based ones. Even though devaluation in the real exchange rate can boost Mexican exports, it will also make component imports more expensive. Considering that Mexico-based exporters have limited power to pass through increased costs onto prices, it will change the mark-up to keep the volume more stable.

Brazil-based producers are not so engaged in exports as Mexico-based producers are. Therefore, the necessity of maintaining stable export flows is not so vital in comparison to that imperative for Mexico. A stable mark-up is more favorable than a stable export volume, leading to a lower variance for the exchange rate coefficient equation.

Figure 3 presents the real exchange rate coefficient for the Brazilian equation. Figure 4 shows the Mexican equation, with a 95% confidence interval (dotted lines). The real exchange rate appears in equation (1) in the natural logarithm. The θ_t^2 coefficient can be interpreted as elasticity: the component imports will rise θ_t^2% given a 1% devaluation in the real exchange rate.

Figure 3: Real Exchange Rate Coefficient for Brazilian Equation.

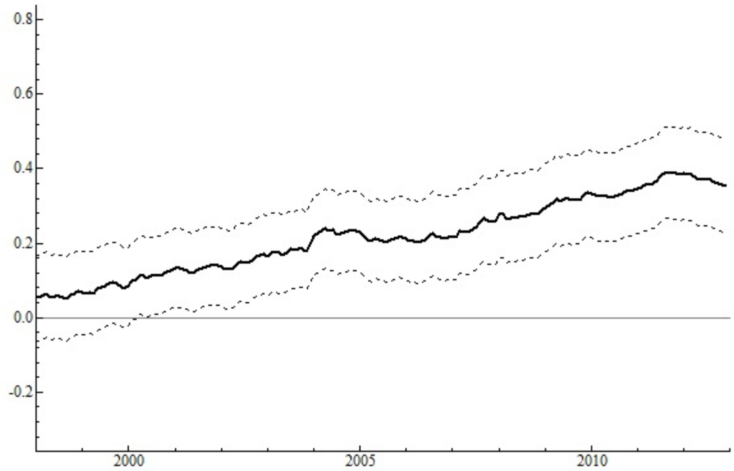
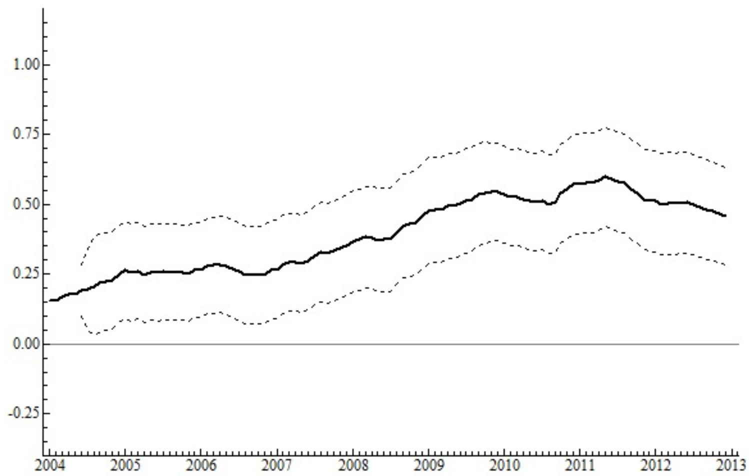


Figure 4: Real Exchange Rate Coefficient for Mexican Equation.



The negative effect that nominal exchange rates have on imports is apparent only in the Brazil equation. Its estimation is presented in Table 6. This is other evidence demonstrating how Brazil is regarded as riskier than Mexico. When the expected nominal exchange rate volatility increases by one basis point, imports tend to drop 0.23%. Because import decisions are regarded as riskier in Brazil than in Mexico, producers in Brazil tend to import more components when exchange rate volatility is low and store them for use during periods of higher volatility. This behavior is completely different in Mexico, where exchange rate volatility does not impact imports, and changes how Japanese companies implement the "just in time" production framework in Brazil as an attempt to mitigate higher risk.

Table 6: Estimated Fixed Effects

Variable	Brazil		Mexico	
	Estimated Coefficient	**P-Value**	**Estimated Coefficient**	**P-Value**
ER Volatility t-1	-0.23	0.03	-0.15	0.71
ER Volatility t-12	-0.09	0.56	0.07	0.87
Ln RER t-1	0.04	0.83	0.64	0.05
Ln RER T-12	Set as variable through time		Set as variable through time	
Ln Final Goods Export to world t-1	0.11	0.25	0.14	0.22
Ln Final Goods Export to world t-12	Set as variable through time		0.1	0.38

With the base-estimated base model, testing for endogenous structural breaks can be done following Harvey and Koopman (1992). A structural break can be regarded as a permanent shock or a shock that takes much time to be dispersed. According to the applied methodology, each country has three dates that represent structural breaks. Only July 2008 shows up in both countries. July 2008 is set as the beginning of the United States financial crisis. For Brazil, the estimated impact is a 27.4% drop in imports from Japan and a 20.5% drop in Mexican imports. The differences in how each country reacted to the 2008 crisis can support the vision that Brazil is regarded as riskier than Mexico. Because the Mexican automobile industry is more dependent on the United States' economic conditions, an economic downturn in United States can be expected to affect Mexico more strongly than Brazil, but the opposite was true, which constitutes additional evidence that Brazil-based producers are more risk-averse.

Table 7: Estimated Structural Breaks

Time Period	Coefficient	P-Value	Long Run Impact
Brazil			
Oct 2003	-0.16	0.001	-14.79%
July 2008	-0.32	0.005	-27.39%
May 2010	0.27	0.004	31.00%
Mexico			
April 2007	0.21	0.00001	23.37%
July 2008	-0.23	0.0001	-20.55%
March 2011	0.14	0.006	15.03%

7 Conclusion

As Moreira (2007) describes, firms can explore considerable market potential and features in Latin America, especially if considering the recent economic downturn in Europe and the United States. However, political and economic risks might be confronted differently by each Japanese firm according to their respective strategies. One possibility is the application of existing production facilities to address internal markets and continuing fragmentation processes, especially in Brazil. Either way, consideration of the possible gains and actual risks is important to reach a better understating of how Japanese firms are managing the fragmentation process and the triangular trade pattern that emerges with it in Latin America.

This study was conducted to generate empirical evidence related to triangular trade patterns of the automobile sector between Japan and the two large Latin America economies, Brazil and Mexico, using state space methodology, and considering a monthly dataset for January 1997 - December 2012.

The results suggest that the effect of real exchange rates is not constant for the countries. As the real exchange rate decreases, exports from Brazil and Mexico tend to grow, which impacts the demand for components and which explains the positive coefficient. Because the one year lag is the only effect found to be significantly different from zero, it apparently takes some time for real exchange rates to affect this trade pattern with Japan.

From the final goods exporter perspective and in its local currency, a low real exchange rate might boost exports. However, those exports become more expensive because of higher prices of imported components. Because firms in the automobile sector can determine their final prices, the exporter might change the markup to compensate for the effects of increasing costs. Because the impact of the real exchange rate on imports is set to change over time, its variation might be interpreted as eagerness of Mexican and Brazilian firms to pass exchange rate variation through into prices.

Observing the graph for the real exchange rate coefficient, the Mexican component is higher than the Brazilian one. Moreover, it is more volatile. This fact might help to describe how Japanese automobile manufacturers regard each country strategically. Considering that most Mexican automobile production is exported, Mexico-based firms cannot always wait for the most favorable exchange rate. Even if hedged, firms are more eager to adjust the markup, leading to a more volatile exchange rate coefficient.

From the beginning of the sample, the real exchange rate coefficient for Brazil is close to zero. It grows gradually over the time. If Brazilian factories are built to supply only the Brazilian market, then there is little incentive to increase exports given devaluation of the real exchange rate. Later, as fragmentation in Brazil progresses and Japanese firms begin to export strategically from Brazil, markups and prices become more important factors than the real exchange rate.

Nominal exchange rate volatility also exerts important effects on triangular trade pattern dynamics. The more volatile the nominal exchange rate is, the riskier or more costly import\export activities become, thereby negatively affecting profitability through higher uncertainty related to costs from imported components. To interpret the estimated impacts that nominal exchange rate volatility has on trade structure, results suggest that real-yen and peso-yen rates are not correlated with Brazilian and Mexican largest automobile market currencies, so the yen volatility will impact only the import decision.

The impact of the exchange rate volatility for the Mexican triangular trade structure is statistically equal to zero. For the Brazilian triangular trade structure, it is estimated as -0.23: a 0.23% drop in imports occurs from each basis point of increased volatility. These differences can be interpreted as evidence of how Japanese companies behave differently in both countries.

Japanese automobile companies established in Mexico have the main objective of exporting to the United States. This trade flow is an important incentive for component imports from Japan, resulting in a steady demand for components from Japan. Import decisions are made with little influence of exchange rate volatility and the risk levels that it represents.

Evidence that exchange rate volatility affects component imports from Japan to Brazil might be related to the fact that Japanese companies in Brazil emphasize supplies to local markets rather than exports. Because Brazilian demand for automobiles is influenced by uncertainty represented by macroeconomic variables such as exchange rates, Japanese companies are more risk-averse. During periods with higher exchange rate volatility, import decisions are postponed. When volatility drops, Japanese companies tend to import aggressively to build up inventories that are useful during high-volatility periods. The necessity to accumulate components engenders higher production costs and lower competitiveness, which are factors that strengthen risk-averse behavior with Japanese companies.

The structural break result shows periods with long-term shocks for each country. Three endogenous structural breaks were estimated for each country. An interesting point is that July 2008 represents a structural break for both countries, with a sharp fall effect on imports (27.4% for Brazil, 20.5% for Mexico). That time represents the beginning of the 2008 United States banking crisis and the ensuing recession. The shock of the 2008 recession more strongly affected Brazil than Mexico, which might also represent how Brazilian producers are more risk-averse than Mexican ones.

Other results that underscore the differences of the two countries are the final goods exports coefficients. For Brazil, it is positive and growing over time. For Mexico it is statistically equal to zero. The Brazilian case might be interpreted as evidence of intensification of the fragmentation process and its development as an export base, which differ completely from the Mexican reality because its exports to the United States constitute the major component of Japanese firms' strategy. With Mexican export growth, Japanese firms might be expected to intensify the fragmentation process by developing component production in Mexico rather than importing components. Consequently, Mexican production and exports are likely to be increasingly independent from imported components.

Evidence suggests great differences in the respective structures of both countries, with Mexico advancing faster than Brazil, because the fragmentation process is occurring with components and not with final automobiles. Aside from Brazil's great market potential, the process is not advancing as rapidly as it might.

Reasons might include fiscal inefficacy, political issues, and lack of infrastructure, which limit gains from trade and openness.

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