

CHINESE ECONOMIC GROWTH AND VISITORS TO JAPAN: A BIVARIATE COINTEGRATION ANALYSIS



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ABSTRACT

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In recent years, Japan has experienced an unprecedented boom in foreign visitors, in particular from mainland China. Much of this boom has been driven by rising living standards in China. This paper studies the long-run relationship between Chinese economic growth and the number of visitors to Japan, using cointegration analysis. The Engle-Granger and the Phillips-Peron tests confirm that these variables are cointegrated. A cointegrating equation and an error correction model are estimated to analyse the recent development of Chinese tourism demand and makes forecasts in the next several years. The cointegrating equation predicts that about 4.2 million Chinese travellers will visit Japan in 2020, up 64.0 percent from 2014. The error correction model indicates that about half of the recent surge of Chinese visitors is attributed to unexpected shocks and not a long-run trend.

Contribution/ Originality: This study is one of very few studies which have investigated international travel demand to Japan with statistically sound methods to analyze time series.

1. INTRODUCTION

In the past few years, Japan has been enjoying an unprecedented boom in foreign visitors. Much of this boom has been driven by the surge of visitors from mainland China. In 2014, for example, more than 2.5 million Chinese landed in Japan, up 58.1 percent from the previous year. In the year, mainland China was the third largest source of foreign visitors, only after South Korea (3.0 million) and Taiwan (2.8 million), out of the total of 14.2 million visitors, according to the [Ministry of Justice of Japan \(2015\)](#).

The number of arrivals from mainland China to Japan expanded at an average rate of more than 10 percent per year from less than 200 thousand arrivals in the early 1990s and surpassed 1.7 million in 2010 (Figure 1). Although it dropped to about 1.3 million after the earthquake and tsunami in 2011, Chinese travellers have returned to Japan with 2012 and 2014 breaking all-time records.

In the long run, economic growth and improvement in living standards, along with political liberalization and globalization, have expanded opportunities for Chinese to travel overseas (see the studies in [Arlt and Burns \(2013\)](#)) for the recent development of Chinese outbound tourism). Chinese real GDP per capita increased roughly a

sevenfold from the early 1990s to 2014 (see Figure 2 from the International Monetary Fund (2015)). According to World Tourism Organization (2015) the number of departures from China for international tourism increased more than 20 times during this period of time, which implies that the long-run income elasticity of Chinese tourists traveling overseas is approximately three.

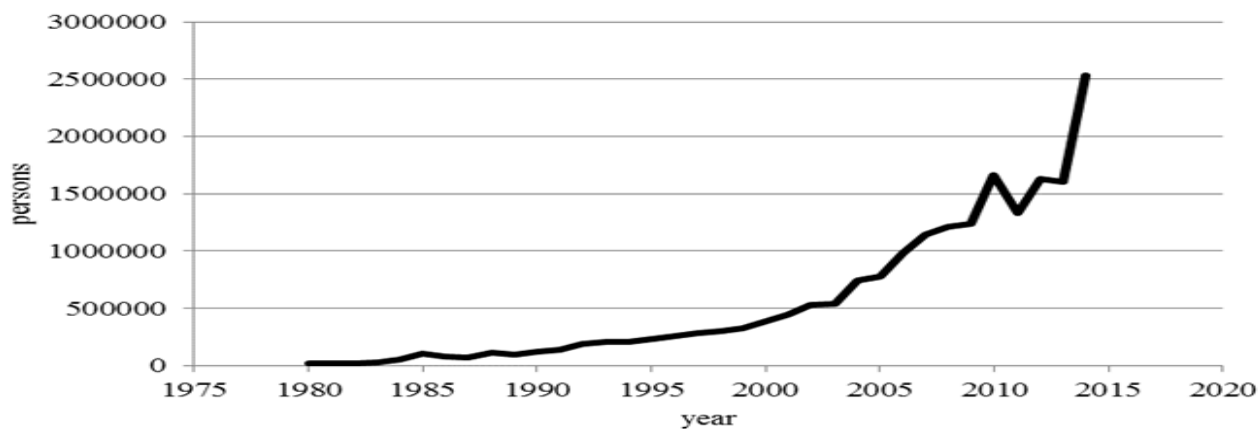


Figure-1. The Number of Arrivals from Mainland China to Japan

Source: The MJJ (2015)

The tourism from China to Japan has been on a growth path over the last decades and the underlying upward trend will continue in the foreseeable future since Chinese economy is expected to grow at a moderate, if not spectacular, rate (see IMF's estimates on Figure 2) and the income elasticity is assumedly positive. It is, however, not certain whether the rapid increase seen in the last few years will be sustainable. There are some factors that seem to have helped fuel Chinese tourism in Japan in the short run. The Japanese government has targeted 20 million overseas visitors by the year of the 2020 Olympics Games in Tokyo and relaxed related visa rules for several countries, including China. A weaker yen is also said to have made traveling and shopping in Japan cheaper. In addition, a fraction of the recent surge of Chinese visitors might have been a recovery to an underlying trend from the sharp drop after the 2011 earthquake. After all these temporary effects decay, it is highly probable that the pace will decelerate to a sustainable path in years to come.

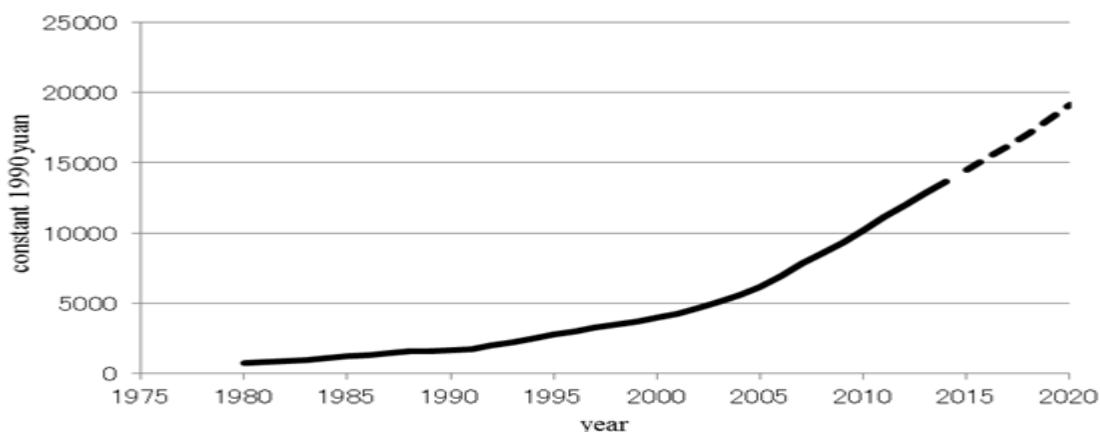


Figure-2. Chinese Real GDP per Capita (Estimates after 2014)

Source: The IMF (2015)

In all planning activities, accurate forecasts of expected demands are crucial. Modeling and forecasting Chinese visitors are, therefore, of considerable importance and interest to Japanese businesses and policy makers, who make decisions, such as recruitment and capital expenditures, on the basis of long-term perspectives. With the volatile inflows of Chinese visitors in the recent years, however, forecasting errors from simple extrapolation will not be

negligible. The long-term and stable demand relationship must be adequately identified and estimated for optimal decision-making.

Demand for international tourism, inbound or outbound, has been extensively studied in the literature of tourism economics and econometrics. In particular, the continuing expansion of worldwide travels since the 1990s has stimulated academic and practical interests in econometric analysis (see [Li et al. \(2005\)](#) and [Song and Li \(2008\)](#) for the survey). In the literature, however, few econometric studies on tourism demand to Japan from China and other origin countries are found. One of a few exceptions is a recent study by [Asemota and Bala \(2012\)](#).

This paper describes the long-run relationship between Chinese economic growth and visitors to Japan and forecasts the expected long-term trend, using cointegration analysis. Many time series variables individually wander apart and follow random walks, but some of them are bounded by equilibrium relationship. These variables are said to be cointegrated. Loosely speaking, variables are cointegrated if non-stationary variables (of the same degree of integration) move together over time. Although cointegrating variables temporarily drift away from one another, their equilibrium relationship will be restored over time. Cointegration analysis has been applied to international tourism demand; inbound tourism demand by [Asemota and Bala \(2012\)](#) for Japan, [Dritsakis \(2004\)](#) for Greece, [Witt et al. \(2004\)](#) for Denmark, [Song and Witt \(2003\)](#) for Korea, [Vogt and Wittayakorn \(1998\)](#) for Thailand, and [Syriopoulos and Sinclair \(1993\)](#) for Mediterranean countries, and outbound demand by [Song et al. \(2000\)](#) for the United Kingdom.

As statistically tested later in this paper, Chinese real GDP per capita and the number of arrivals from mainland China to Japan are cointegrated. Thus, the cointegrating equation, which has the number of arrivals as a dependent variable and real per-capita GDP as an independent variable, is not spurious and will be of great value for forecasting the trend of Chinese visitors under a projected path of real GDP. Applying cointegration analysis, this paper examines the long-run equilibrium and predicts the trajectory of Chinese visitors to Japan conditional on growth forecasts.

The rest of the paper is organized as follows. The next section presents a simple framework modeling cointegrating relationship and correction mechanism from short-run deviation to long-run equilibrium. A model for the error correction mechanism is called an error correction model (ECM), which is transformed from a cointegrating equation. Section 3 describes the data and the results of unit root and cointegration tests. Section 4 presents the empirical results of the cointegrating equation and the ECM. It also extends analysis to the effect of travel cost, which has been extensively investigated in the literature of international tourism demand. It turns out that the cost variable is not of use in forecasting Chinese visitors to Japan in the long run. Finally, forecasts are made for a projected path of Chinese economic growth. The last section concludes.

2. COINTEGRATION AND THE ERROR CORRECTION MODEL (ECM)

Under the assumption that two variables are cointegrated, an ECM is derived from the cointegrating relationship. Following [Cottrell \(2004\)](#) this section describes how the ECM is connected to cointegration, and the models estimated in the sections below.

An ECM generally starts from a long-run equilibrium relationship between two variables:

$$Y_t = K + (X_t)^\gamma \quad (1)$$

In this paper, Y_t is the number of arrivals from China to Japan at time t and X_t is Chinese real income per capita at t . γ represents the income elasticity of visitors, and is also called a cointegrating coefficient in the statistical literature. K is a fixed coefficient. Taking the log of the equation (1), this equation can be written as

$$y_t = k + \gamma x_t \quad (2)$$

The lower-case letters are the logs of the variables and the coefficient.

In the short run, the dependent variable y_t deviates from its long-run value before it reverts to the equilibrium over time. A general functional form is assumed to model short-run dynamics:

$$y_t = \beta_0 + \beta_1 x_t + \beta_2 x_{t-1} + \beta_3 y_{t-1} + e_t \quad (3)$$

This equation includes lagged values x_{t-1} and y_{t-1} , and an error term e_t . β_i are fixed coefficients.

In order to relate the long-run equilibrium relationship (2) and the short-run dynamics (3), suppose that the variables are both in equilibrium at t and $t-1$, and no error term disturbs the relationship. Let x^* and y^* denote equilibrium values, and set $e_t = 0$. Then, the equation (3) becomes

$$y^* = \beta_0 + \beta_1 x^* + \beta_2 x^* + \beta_3 y^* \quad (4)$$

Rewriting this yields

$$y^* = \beta_0 + \beta_1 x^* + \beta_2 x^* + \beta_3 y^* \quad (5)$$

The coefficients in (5) corresponds with those in (2):

$$k = \beta_0 / (1 - \beta_3) \quad (6)$$

$$\gamma = (\beta_1 + \beta_2) / (1 - \beta_3)$$

Assuming these, the short-run dynamic equation (3) can be written in deviation form as

$$\Delta y_t = \beta_0 + \beta_1 \Delta x_t + \delta (y_{t-1} - k - \gamma x_{t-1}) + e_t \quad (7)$$

where $\Delta x_t = x_t - x_{t-1}$, $\Delta y_t = y_t - y_{t-1}$ and $\delta = \beta_3 - 1$. This specification is known as the an error correction model, or the ECM, which relates a change in a dependent variable to a change in an independent variable as well as an error correction term $y_{t-1} - k - \gamma x_{t-1}$. The error correction term represents the deviation from the equilibrium relationship (2) at the previous period. δ describes the speed of adjustment back to the equilibrium, and is expected to be negative.

The parameters in the models (2) (with an error term) and (7) are sequentially estimated in Section 4 after testing whether the variables are cointegrated in Section 3.

3. UNIT ROOT AND COINTEGRATION TESTS

3.1. Data

The data for this paper is annual time series for the period 1980-2014. The tourist arrival data is sourced from the Ministry of Justice of Japan. The Ministry of Justice collects Chinese data by region: "China", "China (Taiwan)", "China (Hong Kong)", and "China (Others)". This paper uses "China", which counts the number of arrivals from mainland China. The real GDP per capita data is obtained from IMF's World Economic Outlook Database October 2015, which also includes forecasts from 2015 to 2020. IMF's forecasts are used for a projected path of Chinese income growth. The number of arrivals is adjusted for the size of population, that is, y_t is the number of arrivals per million. The population data are also sourced from the World Economic Outlook Database.

3.2. Unit Root Test

Before testing whether the two variables are cointegrated, the order of integration of the variables must be determined by unit root test since cointegration analysis requires that variables in a cointegrating system are of the same order of integration. Roughly speaking, a variable is integrated of order 1 if its original series has a unit root while its first difference does not (higher orders of integration are not considered here). In this paper, the augmented Dickey-Fuller (ADF) test is carried out for unit roots on both Chinese real income per capita x_t and the number of Chinese arrivals to Japan y_t per million [Dickey and Fuller \(1979\)](#). For x_t , the ADF test is based on the regression

$$\Delta x_t = \alpha_0 + \alpha_1 t + \alpha_2 x_t + \sum_{i=1}^l \varphi_i \Delta x_{t-i} + u_t \quad (8)$$

where t is a deterministic time trend, u_t is an error term, and α_0 , α_1 and φ_i are all fixed coefficients. The lagged values of Δx_{t-i} are included to correct autocorrelation in u_t . The lag length l is determined by information criterions, such as the Akaike Information Criterion (AIC) and the Bayesian Information Criterion (BIC) [Akaike \(1973\)](#) and [Schwarz \(1978\)](#). The time trend t must be included in the regression (8) not to reduce the power of the test if it is statistically significant.

The null and alternative hypotheses are

$$\begin{aligned} H_0: \alpha_2 &= 0 \\ H_1: \alpha_2 &< 0 \end{aligned} \tag{9}$$

The null hypothesis implies that the variable x_t has a unit root, that is, it is non-stationary. The alternative hypothesis is that the variable does not have a unit root and is stationary along the deterministic trend. Unless the null hypothesis is rejected, the test is sequentially performed for the first difference $\Delta x_{t-1} = x_t - x_{t-1}$. Similarly, y_t and $\Delta y_{t-1} = y_t - y_{t-1}$ are also tested for unit roots.

Table 1 reports the results from the ADF tests on the levels (x_t and y_t), and the first differences (Δx_{t-1} and Δy_{t-1}). In all cases, the AIC and the BIC select identical lag lengths. Since the time trends are not statistically significant for the first differences, the ADF statistics without time trends are also computed. The ADF statistics are compared with the critical values at the 1% significance level. The null hypotheses are not rejected for the levels, but rejected for the first differences, concluding that both of them are integrated of order 1, or I(1).

Table-1. The Augmented Dickey-Fuller Tests for Unit Roots

variable	ADF statistic	lag length
x	-3.829	1
y	-3.567	1
Δx (with trend)	-4.283	1
Δy (with trend)	-4.928	1
Δx (without trend)	-4.377	1
Δy (without trend)	-4.729	1

Source: Author

Notice that the null hypotheses are rejected for the levels if the significance level is set at the 5% level. This might indicate that the variables are both integrated of order 0, or I(0). If so, the equation (2) is safely estimated since a linear combination of I(0) variables is generally I(0) and regressing y_t on x_t is not spurious. Whether x_t and y_t are I(0) or I(1), however, the variables are of the same order of integration. The following analysis only requires that variables are all integrated of the same order.

3.3. Cointegration Test

Given that x_t and y_t are integrated of the first order, they are tested for cointegration to see if there is a stable linear relationship between them in the long run. With an error term v_t , suppose that the long-run relationship (2) holds:

$$y_t = k + \gamma x_t + v_t \tag{2}'$$

The error term v_t should not have a unit root if x_t and y_t are cointegrated, but it should if they are not. Thus, a test for cointegration is based on the regression

$$\Delta v_t = \alpha_2 v_t + \sum_{i=1}^l \varphi_i \Delta v_{t-i} + u_t \tag{8}'$$

and the hypotheses (9). The null hypothesis implies that the variables x_t and y_t are not cointegrated. The ADF statistic can be used, but the sampling distribution of the test statistic differs and a new set of critical values must be calculated because the regression (8)' is on the residuals of the estimated regression, not raw data. Engle and Granger (1987) and Engle and Yoo (1987) tabulate the critical values for the residual-based ADF test, which are larger than those for raw data. This test is known as Engle-Granger (EG) test.

A challenge that the EG test poses is that the hypothesis testing is highly sensitive to the choice of the lag length. In addition, the poor performance of standard information criteria, such as the AIC and the BIC, in selecting the lag length is well known in the statistical literature Sargan and Bhargava (1983); Nishii (1988); Cheung and Lai

(1993); Lütkepohl (1993); Gonzalo (1994); Bewley and Yang (1998); Clarke and Mirza (2006). In this paper, the ADF statistics are computed with several different values of the lag length.

Table 2 shows the results from the EG test. The residuals are calculated from the estimated regression (2) reported on Table 4. The lag length varies from 0 to 4. The null hypothesis is rejected at the 1% significance level if the lag length is 3 or 4. This concludes that the variables are cointegrated although shorter lag lengths tilt the conclusion toward the acceptance of the null hypothesis.

Table-2. The Engle-Granger Tests for Cointegration

ADF statistic	lag length
-2.884	0
-3.382	1
.3.997	2
-5.162	3
-4.737	4

Source: Author

With the somewhat ambiguous conclusion above, the Phillips-Perron test is also carried out to reassure the cointegration (Phillips and Perron, 1988). The results are reported on Table 3 (technical details are not discussed here). It confirms that the variables are cointegrated at the 1% significance level.

Table-3. The Phillips-Perron Test for Contergration

statistic	lag length
-9.693	3

Source: Author

4. EMPIRICAL RESULTS

4.1. The Cointegrating Equation and the ECM

Given that the variables x_t and y_t are cointegrated, the parameters in the models (2) and (7) can be estimated. There are three commonly used methods applicable to the cointegrating equation and the ECM: Engle and Granger (1987); Engle and Yoo (1987) and Johansen (1991) methods. In this paper, the Engle-Granger method is used for simplicity. One drawback of this method is that it is not possible to perform any hypothesis tests on the cointegrating coefficient γ Engle and Granger (1987). In what follows, the estimate of γ is simply taken as a point estimate and not tested for any hypotheses.

The Engle-Granger method takes two steps. In the first step, the cointegrating regression (2)' is estimated by the ordinary least squares (OLS). The residuals are saved for the error correction term $y_{t-1} - k - \gamma x_{t-1}$. In the second step, the ECM is estimated by the OLS with the residuals. It is valid to perform statistical tests on the coefficients in the ECM.

Table 4 reports the results for the cointegrating equation (2). The estimate for the income elasticity Υ is 1.449. That is, a 1 percent increase in real GDP per capita leads to a 1.449 percent increase in the number of arrivals per year. This estimate is line with the other countries in Asemota and Bala (2012) which estimate the income elasticity of arrivals to Japan from Australia (3.317), Canada (2.815), Germany (2.198), the United Kingdom (1.158) and the United States (0.775) (the numbers in the parentheses are the income elasticity for each country). Since the Durbin-Watson statistic indicates positive autocorrelation, the regression with an autocorrelated error term is also estimated by the generalized least squares (the results are not reported here). This yields a similar, but slightly higher estimate of 1.551.

Table-4. OLS Estimates for the Cointegrating Equation

		estimate
constant		-6.301
real GDP per capita		1.449
R squared	0.965	
Durbin-Wats on statistic	0.627	

Source: Author

Table 5 reports the results for the ECM. The estimate on the speed of adjustment δ has the expected negative sign, and implies that the short-term disequilibrium at the previous period is corrected about 60 percent at the current period (40% of the disequilibrium remains after one year). The estimate of β_1 , which represents the short-run income elasticity, is slightly larger than the long-run elasticity γ . The Durbin-Watson statistic indicates no autocorrelation.

Table-5. OLS Estimates for the Error Correction Model

	estimate	standard error
Δx	1.778	0.330
error correction term	-0.397	
R squared	0.542	
Durbin-Wats on statistic	1.883	

Source: Author

4.2. The Effect of Travel Cost

In the recent boom of inbound tourism from China in Japan, it is said, particularly in the media, that a depreciated yen has lowered the costs of traveling and shopping in Japan and attracted Chinese visitors. This section examines the effect of the costs.

Microeconomic theory generally assumes that a demand is a function of its own price. In empirical studies on international tourism demand, thus, some measures of travel cost are also examined as independent variables. In practice, however, it is difficult to obtain those measures, such as transportation and living cost, in officially released data. Instead, the price index of a destination country relative to an origin country is used as a proxy [Asemota and Bala \(2012\)](#). The relative price is expected to influence potential travelers' decisions on whether they travel domestically or internationally.

Figure 3 plots the relative price index between China and Japan from 1980 to 2014. The index is calculated as

$$\frac{(P_J/E_J)}{(P_C/E_C)} \quad (10)$$

where P_J and P_C are the price indexes in Japan and China, and E_J and E_C are the foreign exchange rates expressed in their national currencies per a unit of US dollar. GDP deflators are used for P_J and P_C , and the data are sourced from IMF's Economic Outlook Database October 2015 (2015). Japanese Yen and Chinese Yuan per US dollar are used for E_J and E_C , and the data are obtained from FRED of the [Federal Reserve Bank of St. Louis \(2015\)](#). In FRED, the data series of the Chinese foreign exchange rate starts from 1981; the index is calculated for the years 1981-2014. The index is normalized as 1 in 1981.

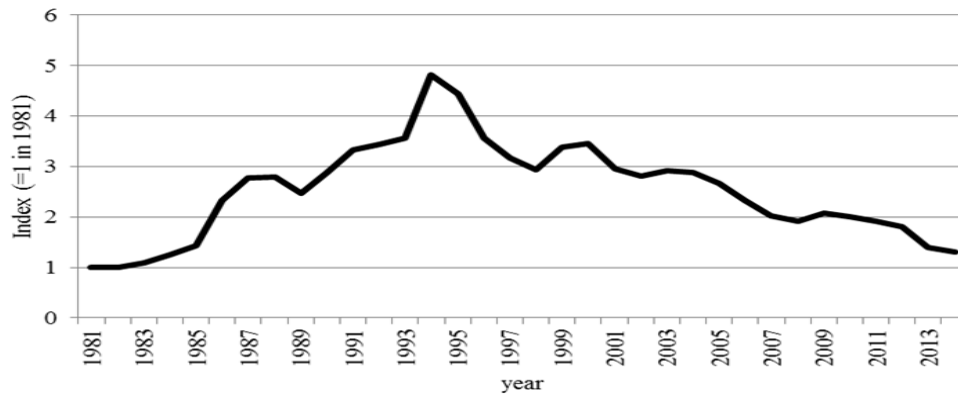


Figure-3. The Relative Price Index: Japan to China

Source: Author

The figure shows that the Japanese price level relative to China rose until the mid-1990s and has fallen over the two decades since then. Considering the fact that the number of Chinese tourists visiting Japan has increased during the entire period of the data, the relative price does not appear to have a unidirectional effect on the Chinese tourism demand in the long run. Also, notice that the decrease in travel and shopping costs in Japan is not a recent phenomenon, and it might not have been an influencing factor behind the surge of Chinese visitors to Japan in the last few years.

Tables 6 reports the results from the multivariate cointegration models including the relative price index. The variables are all integrated of the first order and cointegrated (the results are not reported here). The estimates of the price elasticity have the unexpected positive signs, but are not quantitatively large. The estimates of the other independent variables remain relatively unchanged even after the new independent variable is added to the models. With the unexpected signs of the price elasticity and the stability of the income elasticity, thus, the travel cost variable is to be omitted from the models. In what follows, forecasts are made from the bivariate cointegration system.

Table-6. The Multivariate Cointegration Analysis with the Relative Price Index

A. OLS Estimates for the Cointegration Equation		
	estimate	
Constant	-6.248	
Δx (real GDP per capita)	1.412	
Δx (relative price index)	0.312	
R squared	0.976	
Durbin- Watson statistic	1.110	
B. OLS Estimates for the Error Cointegration Model		
	estimate	standard error
Δx (real GDP per capita)	1.686	0.334
Δx (relative price index)	0.318	0.231
error correction term	-0.638	0.193
R squared	0.564	
Durbin- Watson statistic	1.68	

Source: Author

4.3. Forecasts and Residual Analysis

Figure 4 plots the predicted values of the number of Chinese arrivals in long-run equilibrium, conditional on IMF's growth forecasts from 2015 to 2020 (the predicted values are also multiplied by IMF's forecasts of population). In equilibrium, around 4.2 million Chinese are predicted to visit Japan in 2020, up 64.0 percent from 2014. This is equivalent to the average compounding growth rate of 8.6 percent per year.

The estimated models can be used to quantitatively evaluate unexpected shocks and disequilibrium in the data. The realizations of the error term e_t in the ECM (7) are considered to be unexpected shocks to Chinese tourism demand to Japan. Figure 5 plots the predicted values of the unexpected shocks (the predicted values are also multiplied by population). Large negative shocks are observed in the years 2007-2009 during the Global Financial Crisis and in the years 2011-2013 after the earthquake and tsunami in 2011.

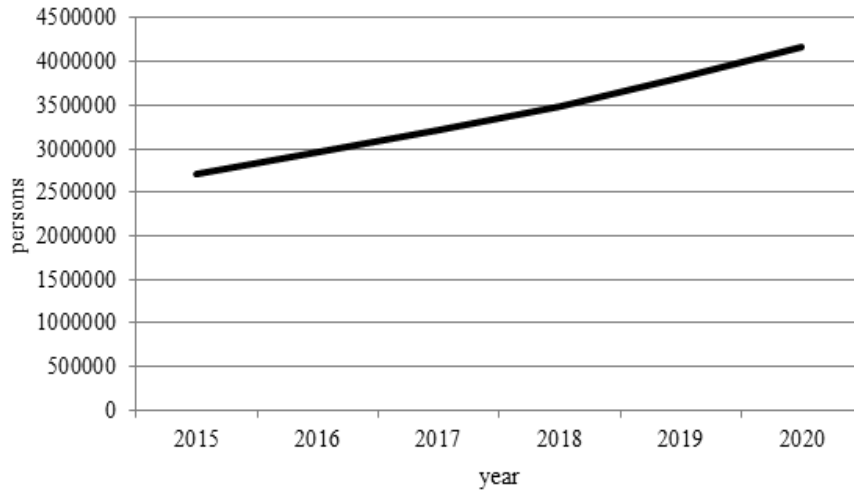


Figure-4. The Predicted Values Number of Arrivals in Equilibrium

Source: Author

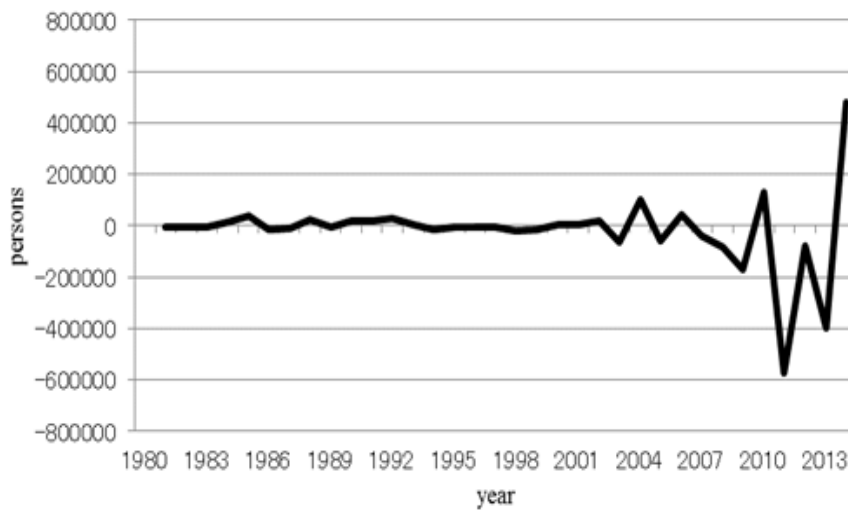


Figure-5. The Predicated Values of the Unexpected Shocks

Source: Author

Figure 6 plots the predicted values of the error correction term $y_{t-1} - k - \gamma x_{t-1}$, or the residuals in the cointegrating equation (2). These represent the sizes of disequilibrium. In the years 2011-2013, the number of Chinese visitors was far below the long-run equilibrium, but it reverted to the equilibrium in 2014 as the negative disequilibrium was corrected after the earthquake and tsunami and a positive unexpected shock hit the Japanese tourism industry in the year.

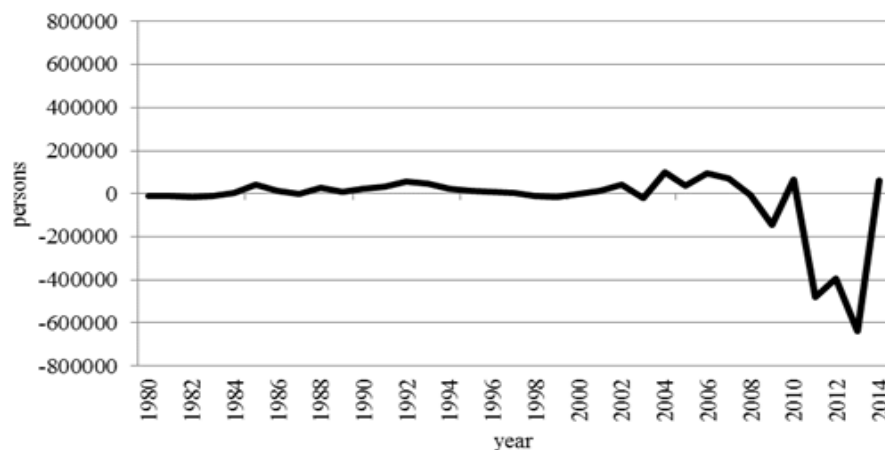


Figure-6. The Predicted Values of the Error Term (Disequilibrium)

Source: Author

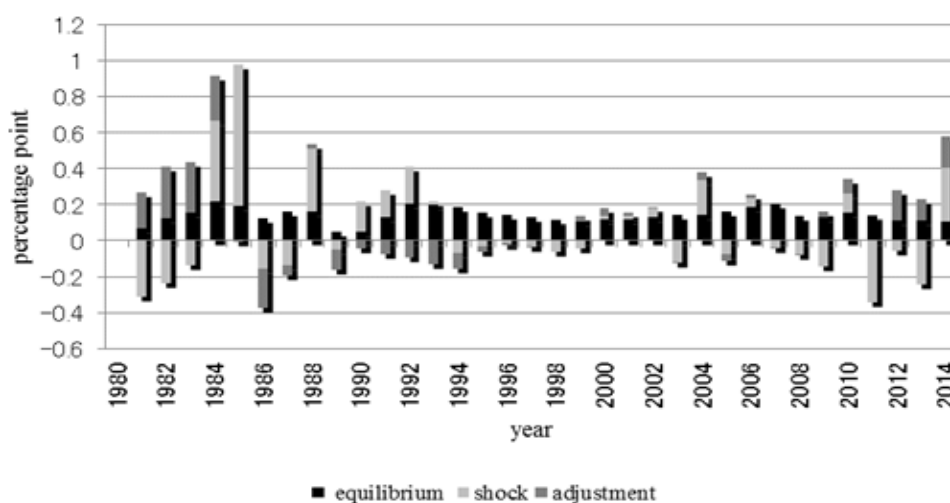


Figure-7. The Decomposition of the Influencing Factors

Source: Author

With all these results, the contributions of the influencing factors to the recent surge of Chinese visitors can be quantified. From 2013 to 2014, the number of arrivals increased about 58.1 percent (calculated from Table 4). Out of this, the contribution of the equilibrium growth was 10.5 percentage points. The unexpected shock boosted the number by 29.8 percentage points (Figure 4). The remaining 17.7 percentage points were due to the slow adjustment of the short-term deviations from the previous year (Figure 5). Figure 6 shows these decompositions for the entire sample period from 1981 to 2014.

5. CONCLUSION

A cointegration equation represents long-run equilibrium relationship between non-stationary but cointegrated variables while an ECM describes short-run adjustment mechanism. This paper has tested the cointegration between Chinese real income per capita and the number of visitors to Japan. The estimated models have allowed the analyses on how much Chinese economic growth has boosted their international travel demand to Japan and unexpected shocks have discouraged potential Chinese travelers to choose Japan as a destination, particularly after the earthquake and tsunami in 2011. The long-run income elasticity of Chinese visitors has been estimated to be around 1.5. With this estimate, this study has predicted that about 4.2 million Chinese will visit Japan in 2020, up 64.0 percent from 2014. It is, however, to be noted that this predicted value is the number of visitors in equilibrium; the expected large inflow of international travelers coming to the 2020 Summer Olympics in Tokyo is not counted

in the models. This paper has also quantified what proportion of the recent surge of Chinese travelers is attributed to long-run equilibrium, unexpected shocks and slow adjustment of disequilibrium. More than half of this is attributed to unexpected shocks, which implies that the rapid increase will decelerate to an equilibrium path after a few years of adjustment.

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