research paper series

China and the World Economy

Research Paper 2011/03

Do Imports Crowd Out Domestic Consumption?
A Comparative Study of China, Japan and Korea

By
Chuanglian Chen, Guojin Chen and Shujie Yao

The Centre acknowledges financial support from The Leverhulme Trust
under Programme Grant F/00 114/AM

Leverhulme Centre
for Research on Globalisation and Economic Policy

The Authors

Chuanglian Chen is a Ph.D. candidate at the School of Economics and Wang Yanan Institute of Economics at Xiamen University, China.

Guojin Chen is professor of Economics at the School of Economics and Wang Yanan Institute of Economics at Xiamen University, China.

Shujie Yao is professor of Economics, School of Contemporary Chinese Studies, University of Nottingham, UK; and special professor of economics, Xi'an Jiaotong University, China.

Acknowledgements

This research is supported by Chinese National Natural Science Foundation (#71071132) and Humanities and Social Sciences Planning Project, Chinese Ministry of Education (#09YJA790118) and Xiamen University Doctorate Research Grant (# ZX11b1).

Do Imports Crowd Out Domestic Consumption?
A Comparative Study of China, Japan and Korea

By
Chuanglian Chen, Guojin Chen and Shujie Yao

Abstract:

A decline in the relative price of imported goods compared to that of domestically produced goods may have different effects on domestic consumption. Such effects may not be accurately detected and measured in a classical permanent-income model without considering consumption habit formation as pointed out by Nishiyama (2005). To resolve this problem, this paper employs an extended permanent-income model which encompasses consumption habit formation. Both cointegration analysis and GMM are used to estimate the (modified) intertemporal elasticities of substitution (IES) between imports and domestic consumption and the parameters of habit formation as well as the (modified) intratemporal elasticities of substitution (AES). We find that import and domestic consumptions are complements in China, but substitutes in Japan and Korea. Different per capita incomes and consumer behaviors between China and the other two countries are two possible reasons for different relationships between import and domestic consumptions. The research findings have important implications on policies such as exchange rate adjustments in China.

JEL: D01, D11, D91
Keywords: Habit Formation; Imports and Domestic Consumption; China, Japan and Korea

Outline

Non-Technical Summary

This paper investigates whether rising imports may crowd out domestic consumption caused by price changes in China, Japan and Korea. When imported goods become relatively cheaper compared to domestically produced goods, caused by domestic currency appreciation or other external shocks, there are two counteractive effects on the demand for domestically produced goods: the intratemporal substitution effect and the intertemporal substitution (or income) effect. Whether imports can crowd out domestic production will depend on the relative forces of these effects.

Such effects may not be accurately detected and measured in a classical permanent-income model without considering consumption habit formation as pointed out by Nishiyama (2005). To resolve this problem, this paper employs an extended permanent-income model which encompasses consumption habit formation. Both cointegration analysis and GMM are used to estimate the (modified) intertemporal elasticities of substitution (IES) between imports and domestic consumption and the parameters of habit formation as well as the (modified) intratemporal elasticities of substitution (AES).

We find that import and domestic consumptions are complements in China, but substitutes in Japan and Korea. Different per capita incomes and consumer behaviors between China and the other two countries are two possible reasons for different relationships between import and domestic consumptions.

As external pressure mounts on China to appreciate its domestic currency, the research results in this paper have direct and relevant policies implications. In the short run, it seems that currency appreciation may not crowd out domestic demand. However, in the long term, as per capita income rises, consumption habit of Chinese consumers may approach that of their Japanese or Korean counterparts. In that case, currency appreciation, leading to a decline in the relative price of imports compared to domestically produced goods, will crowd out domestic demand. Consequent, foreign exchange rate policy reforms have to consider both short and long terms effects in China.

1. Introduction

A decline in the prices of imported goods (imports) has two counteractive effects on the current demand for domestically produced goods (domestic consumption). First, it raises demand for imported goods and crowds out domestic consumption. This is the so-called intratemporal substitution effect. Second, as imported goods become cheaper, real current income rises, leading to higher domestic consumption in the current period at the expense of future consumption. This is the so-called intertemporal substitution, or income, effect.

Whether the intratemporal and intertemporal effects will lead to a net crowding out of domestic consumption will depend on the relative sizes of the intratemporal elasticity of substitution (AES, hereafter for convenience) and the intertemporal elasticity of substitution (IES, hereafter for convenience) of domestic consumption. [1]

If AES is larger than IES, a decline in the prices of imported goods will reduce domestic consumption, or *vice versa*. It is worth noting that a decline in the relative prices of imported goods *vis-a-vis* domestically produced goods can be caused by domestic currency appreciation. As a result, the empirical results from this study will have some useful implications on foreign exchange policy or other price reforms.

Some empirical studies have investigated IES of both imports and domestic consumption in a rational framework based on a Life Cycle / Permanent Income Model (LCPIM). Ceglowski (1991), for example, investigates the role of intertemporal substitution in US import demand using a model of import consumption based on LCPIM, and estimates the

[1] In section 2, we can see that the IES and AES have to be modified based on habit formation. When habit formation is encompassed, we define them as modified IES and modified AES.

intertemporal elasticity for imports to be about 0.8, while the implied relative price elasticity of import consumption to be about 1. These results indicate that import consumption may respond to changes in their intertemporal prices, as well as changes in their price relative to that of domestic substitutes.

Clarida (1994) employs a simple rational-expectation permanent-income model to derive a structural econometric specification of demand for imported consumer goods. He estimates the average long-run price elasticity of import demand to be -0.95 using a cointegrating approach. The average elasticity of import demand with respect to a permanent increase in real spending was 2.15. Amano and Wirjanto (1996) examine the importance of intertemporal substitution in US import consumption using a model of permanent income that allows for random preference shocks and additive separability of a utility function. Using a cointegtation approach, they show that IES for domestic and import consumption were 0.6 and 0.9, respectively. Using the GMM approach, the estimated IES were 1.4 and 4.3, respectively. However, the J-test tends to reject the model which indicates that IES estimated from GMM appears implausible. The empirical results show that IES estimated from intratemporal optimality condition and from Euler equations are hardly equal.

Nishiyama (2005) argues that, the existence of heterogenous agents, the rich and the poor, and habit formation in the economy seem to explain this empirical dilemma. On the other hand, Muellbauer (1988), Eichenbaum, Hansen and Singleton (1988), Ferson and Constantinides (1991), Ogaki and Park (1997) and Croix and Urbain (1998) all find that habit formation helps to account for consumption dynamics and explains why empirical data frequently reject the life cycle hypothesis.

Habit formation is one form of time-non-separability, which means that the level of consumption is easy to be adjusted upward, but difficult to be adjusted downward. Just like the ancient Chinese proverb "it's easier to go from rags to riches than riches to rags". The idea of introducing habit formation into the utility function can date back to Duesenberry (1949). He assumes that utility in each period not only depends on current consumption, but also on past consumption. Therefore, habit formation can measure the change of consumption on the utility, and describe the irreversibleness of consumption.

Croix and Urbain (1998) extend previous work done by Clarida (1994) and Ceglowski (1991) by considering a two-good version of the lifecycle model introducing time-non-separability in household's preferences, and then use quarterly data for USA and France to test the model. With the information contained in the observed stochastic and deterministic trends, they derive a cointegration restriction to estimate curvature parameters of the instantaneous utility function. The remaining parameters are estimated in a second step by GMM. The constancy of different parameters is investigated both in the long run and the short run. Habit formation turns out to be an important factor of import demand, and negligence of habit formation may lead to frequent rejection of the lifecycle hypothesis.

In order to deal with inconsistent IES estimated from intratemporal optimality condition and from Euler equations, Nishiyama (2005) proposes the cross-Euler equation approach as a prescription for this empirical dilemma, and finds that the Euler equation for domestic non-durable goods is mis-specified, while the Euler equation for imported non-durable goods is somehow correctly specified. Croix and Urbain (1998) and Nishiyama (2005) introduce habit formation into the permanent income hypothesis model and find that habit formation turns out to be an important factor for both import and domestic demands.

In this paper, we first extend the classical permanent-income model by introducing habit formation. Our theoretical model will be more realistic and robust to avoid the empirical dilemma described by Nishiyama (2005). If the parameters of habit formation are set to zero, the model degenerates to the classical model employed by Ceglowski (1991), Clarida (1994), Amano and Wirjanto (1996) and Xu (2002).

We then investigate whether import demand crowds out domestic demand in China, Japan and Korea. Following Cooley and Ogaki (1991), a two-step procedure is used. In the first step, a cointegration approach is used to estimate the cointegrating estimators of IES of import and domestic demands. In the second step, the estimated parameters derived from the first step are plugged into an Euler equation, and use GMM to estimate the parameters of habit formation of import and domestic demands.

The empirical results show that import and domestic consumptions are complements in China, but substitutes in Japan and Korea. It suggests that lower per capita incomes and different consumption behavior of Chinese consumers from their Japanese and Korean counterparts may explain this difference.

The rest of this paper is organized as follows. Section 2 describes the theoretical model incorporating habit formation into a classical two-good permanent income model. Section 3 presents the structural econometric methodology, and methods to calculate Marshallian price elasticities, expenditure elasticities, modified IES and modified AES, and then discusses their implications on the relationship between import and domestic demands. Section 4 provides the empirical data used in this paper. Section 5 reports the empirical results and analyzes whether imports crowd out domestic consumption in China, Japan and Korea. Section 6 concludes.

2. Theoretical model

Ceglowski (1991), Clarida (1994), Amano & Wirjanto (1996) and Xu (2002) employ a two-good permanent income model with additively separable preferences to derive a structural econometric equation and then take full advantage of the well-developed theory of cointegration to investigate the relationship between imported and domestically-produced goods. However, there would be an empirical dilemma, as IES parameters estimated from the intratemporal optimality condition and from Euler equations are inconsistent. Nishiyama (2005) argues that the existence of heterogenous agents, the rich and the poor, and habit formation in the economy seem to explain this puzzle.

In order to overcome this problem, we introduce habit formation into the additively separable instantaneous utility function of the representative household. Consumer utility in each period depends on both present and past domestic and import consumptions. Our two-good permanent income model is based on Muellbauer (1988) and Croix and Urbain (1998), where the instantaneous utility function of the representative household is defined as follows.

$$u(D_t^*, F_t^*) = \begin{cases} \dfrac{D_t^{*1-\rho}}{1-\rho} + \dfrac{F_t^{*1-\upsilon}}{1-\upsilon} & if\ \rho \neq \upsilon \neq 1 \\ \ln D_t^* + \ln F_t^* & if\ \rho = \upsilon = 1 \end{cases} \tag{1}$$

Where $D_t^* = (1-\gamma)^{-1}(D_t - \gamma D_{t-1})$ and $F_t^* = (1-\delta)^{-1}(F_t - \delta F_{t-1})$ are the total flows of domestic and import consumptions, respectively. $\gamma\left(\in[-1,1)\right)$ and $\delta\left(\in[-1,1)\right)$ index the importance of habit formation of domestic and import consumptions. If they are positive, the larger the values are, the greater the impact does previous consumption have on current utility. In order to maximize his or her expected lifetime utility under a lifetime budget constraint, a representative agent would choose to smooth consumption over the whole lifetime. If they are negative, indicating that the goods present some durability (Ferson and Constantinides, 1991), in which case previous consumption still contributes to current utility. Note that, we only consider the impact of one-period lagged consumption on current utility. The dynamic optimization problem of a representative household is formulated as follows.

$$\underset{\{D_t^*, F_t^*\}}{Max} \; E_0\left\{\sum_{t=0}^{\infty}\beta^t u(D_t^*, F_t^*)\right\} \tag{2}$$

Where E_0 is an expectation operator based on period zero information, β a subjective discount factor, P_t^F and P_t^D respectively denote prices of imported and domestically-produced goods. Assuming $P_t = P_t^F / P_t^D$, we can derive the lifetime budget constraint of the agent as follows:

$$A_{t+1} + D_t + P_t F_t \le Y_t + (1+r_t)A_t \tag{3}$$

Where A_t is the real assets held by the household at time t, Y_t is the stochastic labor income at time t, r_t stands for real interest rate from period t to $t+1$. Using the lagrangian approach to solve the above optimal problem, we can obtain an intratemporal or static first-order condition and Euler equations:

$$P_t\left(\frac{1}{1-\gamma}\right)\left(\frac{D_t-\gamma D_{t-1}}{1-\gamma}\right)^{-\rho}\left[1-\beta\gamma E_t\left(\frac{D_{t+1}-\gamma D_t}{D_t-\gamma D_{t-1}}\right)^{-\rho}\right]=\left(\frac{1}{1-\delta}\right)\left(\frac{F_t-\delta F_{t-1}}{1-\delta}\right)^{-\upsilon}\left[1-\beta\delta E_t\left(\frac{F_{t+1}-\delta F_t}{F_t-\delta F_{t-1}}\right)^{-\upsilon}\right] \tag{4}$$

$$E_t\left[\beta(1+r_{t+1})\frac{(D_{t+1}-\gamma D_t)^{-\hat{\rho}}-\beta\gamma(D_{t+2}-\gamma D_{t+1})^{-\hat{\rho}}}{(D_t-\gamma D_{t-1})^{-\hat{\rho}}-\beta\gamma(D_{t+1}-\gamma D_t)^{-\hat{\rho}}}-1\right]=0$$
$$E_t\left[\beta(1+r_{t+1})\frac{(F_{t+1}-\delta F_t)^{-\hat{\upsilon}}-\beta\delta(F_{t+2}-\delta F_{t+1})^{-\hat{\upsilon}}}{(F_t-\delta F_{t-1})^{-\hat{\upsilon}}-\beta\delta(F_{t+1}-\delta F_t)^{-\hat{\upsilon}}}\left(\frac{P_t}{P_{t+1}}\right)-1\right]=0 \tag{5}$$

The above model has two advantages. First, it generalizes the classical model of consumer behavior used in Ceglowski (1991) and others to allow for richer dynamics. In particular, under this scheme, as to the existence of habit formation, current import consumption can be substituted for current domestic consumption (intratemporal substitution) or future import consumption (intertemporal substitution). In fact, if the parameters of habit formation are set to zero, the model degenerates to a classical model in Ceglowski (1991).

Secondly, the model is more realistic by introducing habit formation, as it is one form of time non-separable preferences which are found to be important factors considered by socio-psychologists. In our framework, current utility in each period not only depends on current consumption, but also on past consumption. Furthermore, the static first-order condition and Euler equations derived from this model would be more robust to avoid the empirical dilemma described by Nishiyama (2005).

3. Structural econometric equation and methodology

Taking logarithms on both sides of equation (4) and adopting the linear approximation of one-order Taylor's expansion proposed by Muellbauer (1988), we have,

$$\rho\ln D_t-\ln P_t-\upsilon\ln F_t+c=\frac{\upsilon}{f+(1-\delta)/\delta}\Delta\ln F_t-\frac{\rho}{g+(1-\gamma)/\gamma}\Delta\ln D_t$$
$$+\ln\left[1-\beta\gamma E_t\left(\frac{D_{t+1}-\gamma D_t}{D_t-\gamma D_{t-1}}\right)^{-\rho}\right]-\ln\left[1-\beta\delta E_t\left(\frac{F_{t+1}-\delta F_t}{F_t-\delta F_{t-1}}\right)^{-\upsilon}\right]+o(\ln D_t)+o(\ln F_t) \tag{6}$$

Where $c = \ln[(1-\gamma)/(1-\delta)]$, $o(\ln D_t)$ and $o(\ln F_t)$ denote high-level order terms of $\ln D_t$ and $\ln F_t$, respectively. g and f respectively stand for the average $\Delta \ln D_t$ and $\Delta \ln F_t$. $\ln P_t$, $\ln F_t$ and $\ln D_t$ are cointegrated in equation (6), as long as these variables are $I(1)$.[2] In that case, $\Delta \ln D_t$ and $\Delta \ln F_t$ are $I(0)$ and the right hand side variables in (6) are covariance stationary or ingredients of stochastic disturbance.

Based on Engle and Granger (1987)'s two step method, the asymptotic distribution of GMM estimators in the second step are independent of the first step estimators since the estimated $\hat{\rho}$ and $\hat{\upsilon}$ converge faster than the GMM estimators.[3] In analogy to Cooley and Ogaki (1991), our first step takes the right hand side of (6) as disturbance term ε_t with a cointegrating approach to estimate the cointegrating estimators of IES of import and domestic consumptions. Our second step plugs in the estimated values from the first step into an Euler equation (5), and uses GMM to estimate the parameters of habit formation for import and domestic consumptions.

The first step cointegrating relationship is given by

$$\ln D_t = c' + \frac{\upsilon}{\rho} \ln F_t + \frac{1}{\rho} \ln P_t + \varepsilon_t \tag{7}$$

Where $c' = -c/\rho$, ε_t is I(0) with mean zero. $1/\rho$ denotes IES between domestic consumption and imports, υ/ρ stands for their intratemporal elasticity of substitution (AES). All these parameters are used to calculate the Marshallian price elasticity of imported goods and expenditure elasticities of imported and domestically-produced goods.

---

[2] In this paper, data for China, Japan and Korea seem to support this assumption.

[3] The advantages of using a cointegrating approach to estimate the preference parameters of the utility function is pointed out and discussed by Ogaki (1992) and Ogaki and Park (1997).

The Marshallian price elasticity and expenditure elasticity of imported goods are shown below, respectively. [4]

$$\eta_{F,P} = -\frac{1}{\upsilon}\left[1 - \frac{(1-\upsilon)(1-s)}{(s\upsilon/\rho)+(1-s)}\right] \quad \text{and} \quad \eta_{F,(D+PF)} = \frac{\rho}{\upsilon}\left[\frac{1}{s+(\rho/\upsilon)(1-s)}\right] \tag{8}$$

In an additively separable utility function, according to Ogaki (1992) and Nishiyama (2005), the Marshallian expenditure elasticity of domestic goods is given by

$$\eta_{D,(D+PF)} = \left[\frac{\rho}{\upsilon} + s\left(1 - \frac{\rho}{\upsilon}\right)\right]^{-1} \tag{9}$$

Where $s = P_t^D D_t / (P_t^D D_t + P_t^F F_t)$ denotes the share of spending on domestic goods. Thus, the Marshallian expenditure elasticity of domestic goods, in analogy to the Marshallian price elasticity and expenditure elasticity of imported goods, is also time-varying.

In the second step, estimated coefficients obtained from (7) are plugged into an Euler equation (5). GMM is then used to estimate the parameters of habit formation of import and domestic consumptions.

When habit formation is allowed for, the intertemporal choice becomes more complex. Now, the agents recognize the impact of current choices on their future tastes as to the existence of habit formation, which will render $1/\rho$ and $1/\upsilon$ invalid to measure IES of domestic and import consumption (Constantinides, 1990). However, Boldrin, Christiano and Fisher (1995) and Croix and Urbain (1998) construct IES in a deterministic framework, which is modified by habit formation, or defined as modified IES. Adapting their derivation to our case, the modified IES of domestic and import consumption, are given in (10).

[4] Proofs for equations (8) and (9) are available on request, or see Clarida (1994), Croix and Urbain (1998), Ogaki (1992) and Nishiyama (2005).

$$\frac{1}{\tilde{\rho}} = \frac{1}{\rho}(1-\gamma)\frac{1-\gamma\beta/(1+g)^{\rho}}{1+\gamma^2\beta/(1+g)^{\rho+1}} = a\frac{1}{\rho} \quad \text{and} \quad \frac{1}{\tilde{\upsilon}} = \frac{1}{\upsilon}(1-\delta)\frac{1-\delta\beta/(1+f)^{\upsilon}}{1+\delta^2\beta/(1+f)^{\upsilon+1}} = b\frac{1}{\upsilon} \quad (10)$$

Where β is a subjective discount factor, a and b are modified factors, γ and δ denote habit formation of domestic and import consumptions, respectively. g and f respectively stand for the average $\Delta \ln D_t$ and $\Delta \ln F_t$. Note that $\tilde{\upsilon}/\tilde{\rho}$ is the modified AES between import and domestic consumptions.

According to Amano, Ho and Wirjanto (1998) and Nieh and Ho (2006), there are three testable implications on the relationship between import and domestic consumptions.

(a) If $1/\tilde{\rho} > \tilde{\upsilon}/\tilde{\rho}$, import consumption and domestic consumption are complements, under which, the modified IES of domestic consumption is larger than the corresponding modified AES.

(b) If $1/\tilde{\rho} < \tilde{\upsilon}/\tilde{\rho}$, import consumption and domestic consumption are substitutes, under which, the modified IES of domestic consumption is less than the corresponding modified AES.

(c) If $1/\tilde{\rho} = \tilde{\upsilon}/\tilde{\rho}$, import consumption and domestic consumption are independent, or unrelated.

4. Data

This paper uses data from 1994M01 to 2010M04 (196 observations) for China and Japan. Due to missing observations, Korean data only covers the period 1995M01-2010M04 (184 observations). Monthly data are seasonally adjusted.

Monthly data are constructed in constant US dollars for imports of Food and Direct Consumer Goods for Japan, and imports of Consumer Goods for Korea. As direct import consumption goods data for China are unavailable, they are indirectly obtained using information provided by the United Nations Statistics Division. According to the correspondence between Standard International Trade Classification (SITC) Revision 3 and Broad Economic Categories (BEC), data are derived from 19 BEC basic categories. According to the correspondence between BEC with the basic classes of goods in the System of National Accounts (SNA), data are derived for consumption goods, intermediate goods and capital goods in SNA.[5]

Per capita nominal or real values are obtained by dividing the respective total values by total population. All real values are measured in constant 2005 US dollar prices.[6] As data for domestic goods are unavailable, following Clarida (1994), they are constructed by subtracting per capita import consumption from per capita total consumption (DN_t), which is obtained from dividing total retail sales by total population. Thus, per capita real domestic consumption is defined as follows.

$$D_t = \left(DN_t - P_t^F F_t\right) / P_t^D \tag{11}$$

Where DN_t is nominal per capita consumption expenditures, F_t per capita import consumption, P_t^F implicit price index of imported consumer goods and P_t^D producer price index of domestic consumer goods.[7] The relative price P_t is defined as the ratio P_t^F / P_t^D. Real interest rate is defined as the difference between inflation rate and Interbank Offered Rate for

---

[5] United Nations Statistics Division website provides the detail of correspondence between SITC and BEC, BEC and SNA. Website: http://unstats.un.org/unsd/cr/registry/regdnld.asp?Lg=1.

[6] As China has not yet published Import Price Index of consumer goods and Producer Price Index of manufactured products monthly fixed base index, this paper uses China's Import Price Index of consumer goods and Producer Price Index of manufactured products monthly year-on-year index and seasonally adjusted index to construct China's Import Price Index of consumer goods and Producer Price Index of manufactured products monthly fixed base ratio index (with 2005 as the base year).

[7] We use Import Price Index for Japan and Korea, and Import Price Index of consumer goods for China. Producer Price Index for Japan and Korea, and Producer Price Index of manufactured products for China.

China or 1-month government bond yield for Japan and Korea.

All the data are collected from *IMF*, China Custom Statistics, China's Economic Internet Database (CEInet), China's External Trade Indices, The People's Bank of China, Bank of Japan and CEIC Global Database.

5. Empirical results

Summary statistics are reported in Table 1. The share of spending on domestic goods in China is larger than that in Korea, but smaller than that in Japan. China's import consumption and domestic consumption are all lower than Korea's and Japan's. One noticeable difference among the three countries is that China has the highest volatility in domestic consumption. The negative average value of $\ln P_t$ means that import price index is lower than producer price index for China and Japan, while for Korea it has an opposite meaning. Among the three countries, Korea has the highest real interest rate and Japan the lowest.

Table 1 Summary statistics of selected variables

Country	V	Average	Std. Dev.	Minimum	Maximum	Obs
China	$\ln D_t$	10.3	0.70	9.1	11.7	196
	$\ln F_t$	8.1	0.24	7.3	8.6	196
	$\ln P_t$	-0.2	0.29	-0.9	0.2	196
	r_t	2.2%	0.05	-16.7%	9.6%	196
	s	92.4%	0.03	84.1%	97.5%	196
Japan	$\ln D_t$	11.2	0.13	10.9	11.6	196
	$\ln F_t$	8.1	0.15	7.8	8.5	196
	$\ln P_t$	-0.1	0.15	-0.4	0.3	196
	r_t	1.7%	0.01	-1.2%	4.1%	196
	s	96.1%	0.01	91.7%	97.4%	196
Korea	$\ln D_t$	12.5	0.23	11.6	12.9	184
	$\ln F_t$	10.6	0.30	9.6	11.1	184
	$\ln P_t$	0.03	0.11	-0.1	0.3	184
	r_t	3.9%	0.03	-0.7%	9.7%	184
	s	85.4%	0.02	77.5%	91.1%	184

Note: (1) $s = P_t^D D_t / (P_t^D D_t + P_t^F F_t)$ denotes the share of spending on domestic goods.

(2) The unit of import and domestic consumption is US$ million.

Table 2 presents the results of *ADF* and *PP* tests, the critical values for *ADF* and *PP* tests are given by MacKinnon (1996). In both methods, a constant term is included in the level equation but not in the first difference one. Besides, lag order for *ADF* test is selected by the SC criterion, while bandwidth for *PP* test is selected by Newey-West (1994).

The tested results suggest that the null hypothesis of a unit root cannot be rejected at the

5% critical level. The results of *ADF* and *PP* tests suggest that $\ln D_t$, $\ln F_t$ and $\ln P_t$ are *I(1)*.

Table 2 Unit root test results

	V	Levels				1st dfference			
		ADF		*PP*		*ADF*		*PP*	
		Mode	Stat.	Mode	Stat.	Mode	Stat.	Mode	Stat.
China	$\ln D_t$	(C,N,13)	-0.58	(C,N,12)	-0.41	(N,N,12)	-2.95	(N,N,8)	-13.80
	$\ln F_t$	(C,N,13)	0.39	(C,N,5)	0.88	(N,N,12)	-5.67	(N,N,37)	-28.67
	$\ln P_t$	(C,N,3)	-1.60	(C,N,4)	-2.25	(N,N,2)	-12.59	(N,N,12)	-25.62
Japan	$\ln D_t$	(C,N,13)	-1.49	(C,N,8)	0.22	(N,N,12)	-4.13	(N,N,11)	-81.42
	$\ln F_t$	(C,N,11)	-0.18	(C,N,4)	-0.06	(N,N,10)	-9.28	(N,N,37)	-40.08
	$\ln P_t$	(C,N,1)	-1.97	(C,N,0)	-1.30	(N,N,0)	-8.74	(N,N,6)	-8.41
Korea	$\ln D_t$	(C,N,12)	-1.37	(C,N,14)	-2.33	(N,N,11)	-3.14	(N,N,92)	-25.78
	$\ln F_t$	(C,N,2)	-2.72	(C,N,6)	-1.37	(N,N,1)	-13.71	(N,N,25)	-20.15
	$\ln P_t$	(C,N,1)	-2.20	(C,N,3)	-1.81	(N,N,0)	-9.94	(N,N,11)	-9.52
5% critical values			-2.88		-2.88		-1.94		-1.94

Notes: *ADF* test based on (C,T,K), C=constant, T=trend, K=lag order. *PP* test based on (C,T,B), B=bandwidth.

As all the concerned variables are I(1), the full modified ordinary least squares (FMOLS) and dynamic least squares (DOLS) are used to estimate the long-run cointegrating parameters. According to Phillips and Hansen (1990), Hansen (1992, 2002) and Stock and Watson (1993), FMOLS and DOLS estimators possess the same limited distribution as the full information maximum likelihood estimators and hence are asymptotically optimal. Where FMOLS is based on semi-parametric corrections for endogeneity and serial correlation, by increasing leads and lags of the first differences in the regression can also correct endogeneity and serial correlation. Hence, DOLS estimators are superconsistent and the properly rescaled *t* and Wald statistics for hypotheses about estimators have the conventional asymptotic distributions (standard normal and chi squared). The proper rescaling is to multiply the usual t value by $(s/\hat{\lambda})$ and the Wald statictics by $(s/\hat{\lambda})^2$.[8]

Table 3 Cointegtation regression results: DOLS and FMOLS of equation (7)

	Method	Cst.	$\ln P_t$	$\ln F_t$	*ADF*	L_c	*SupF*	*MeanF*	Implied *IES* $1/\hat{\rho}$	Implied *IES* $1/\hat{\upsilon}$
China	*DOLS*	2.869 (1.52)	1.891 (7.08)	0.946 (1.66)	-3.24	——	——	——	1.891	1.999
	FMOLS	2.939 (2.65)	1.983 (8.47)	0.930 (2.77)	-4.38	0.323 [0.14]	2.622 [0.20]	6.205 [0.20]	1.983	2.132
Japan	*DOLS*	4.140 (4.49)	0.291 (3.03)	0.878 (7.70)	-1.80	——	——	——	0.291	0.331
	FMOLS	4.884 (7.84)	0.252 (3.19)	0.786 (10.21)	-2.28	0.684 [0.01]	6.610 [0.04]	14.10 [0.10]	0.252	0.321
Korea	*DOLS*	6.726 (15.43)	0.343 (4.34)	0.540 (13.17)	-2.53	——	——	——	0.343	0.635
	FMOLS	5.981 (6.21)	0.370 (2.03)	0.610 (6.78)	-2.86	1.035 [0.01]	7.164 [0.03]	13.77 [0.11]	0.370	0.606
1% critical values of test for parameter instability						1.03	8.50	18.6		

Note: (1) Numbers in parentheses are t-values, 10%, 5% and 1% critical values are respectively 1.65, 1.96 and 2.58. Numbers in square brackets are p-values. Critical values of L_c, *SupF* and *MeanF* see Hansen (1992, 2002). (2) FMOLS estimates are based on *VAR(1)* prewhitening procedure and Pzrzen kernel. DOLS estimates are based on one lead and one lag of first differences. (3) $1/\hat{\rho}$ and $1/\hat{\upsilon}$ are respectively implied IES of domestic and import consumption based on equation (7). (4) Null hypothesis of ADF test is no cointegration .

Engle and Granger (1987) suggest applying the *ADF* t-test to the residuals in order to test for the null hypothesis of no cointegration. The sixth column in Table 3 gives the results. No drift is included in the test equation for the level residuals. The test results reject the null hypothesis of no cointegration, meaning that $\ln D_t$, $\ln F_t$ and $\ln P_t$ are cointegrated. The L_c statistics

[8] Where s is standard error when using *OLS* to regress equation (7). A consistent estimate of $\hat{\lambda}$ is obtained as follows: $\hat{\varepsilon}_t$ is residuals of *OLS* regression on equation (7), fitting an *AR(p)* process to the residuals, from $\hat{\varepsilon}_t = \rho_1\hat{\varepsilon}_{t-1} + \rho_2\hat{\varepsilon}_{t-2} + \ldots + \rho_p\hat{\varepsilon}_{t-p} + e_t$, where $t = p+1,\ldots,T$, and then use *AIC* to pick the lag length. Given $\hat{\sigma}^2 = \frac{1}{T-p}\sum_{t=p+1}^{T}\hat{e}_t^2$, then we can derive $\hat{\lambda}^2 = \frac{\hat{\sigma}^2}{(1-\hat{\rho}_1-\ldots-\hat{\rho}_p)^2}$.

cannot reject the null hypothesis of variables cointegrated at the 1% critical level based on FMOLS.

Overall, the results presented in Table 3 are encouraging. They show that the estimated parameters for $\ln P_t$ and $\ln F_t$ from the two approaches are statistically significant with *a priori* expected signs. We also find that the estimators are little different from each other obtained from two different estimation methods.

The FMOLS estimates of IES of domestic consumption for China, Japan and Korea are respectively 1.983, 0.252 and 0.370 and the corresponding AES between import and domestic consumption are 0.930, 0.786 and 0.610. The DOLS estimates of IES of domestic consumption are respectively 1.891, 0.291 and 0.343 for China, Japan and Korea and the corresponding AES between import and domestic consumption are 0.946, 0.878 and 0.540.

These estimated cointegration parameters show that China not only has the largest IES, but also the largest AES. The IES of import consumption can be obtained by dividing the IES of domestic consumption by the AES between import and domestic consumption. The results are given in the eleventh column of Table 3. Obviously, China has the largest IES of domestic consumption. In the second stage, the estimated parameters ($1/\hat{\rho}$ and $1/\hat{\upsilon}$) from the cointegration analysis are plugged into an Euler equation (5) and GMM is used to estimate the parameters of habit formation for import and domestic consumptions.

The columns *SupF* and *MeanF* are derived to test for the consistency of parameters with asymptotic critical values provided by Hansen (1992, 2002). The test results cannot reject the null hypothesis of parameters consistency at the 1% level in all regression models.

Hansen (1992, 2002) constructs a test for cointegrating parameters instability on the basis of FMOLS estimation. The *SupF* test is in the spirit of traditional Chow tests. The procedure is as follows. It first calculates a standard Chow F-statistics for a fixed break point t/T, and then considers the sequence of statistics by varying the location of the break. The final statistics is the following sequence. [9]

$$SupF = \sup_{t/T \in [0.15, 0.85]} F_{t/T} \tag{12}$$

SupF statistics sequence is used to test for cointegrating parameters instability in order to see how a policy shock, e.g., exchange rate adjustment, affects estimated results. The test results are given in Figures 1 to 3 for China, Japan and Korea, respectively.

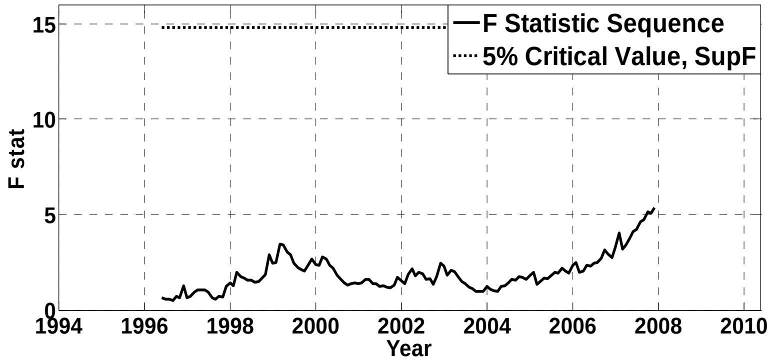

Figure 1 Constancy of the Cointegration Parameters,China: 1994:01-2010:04

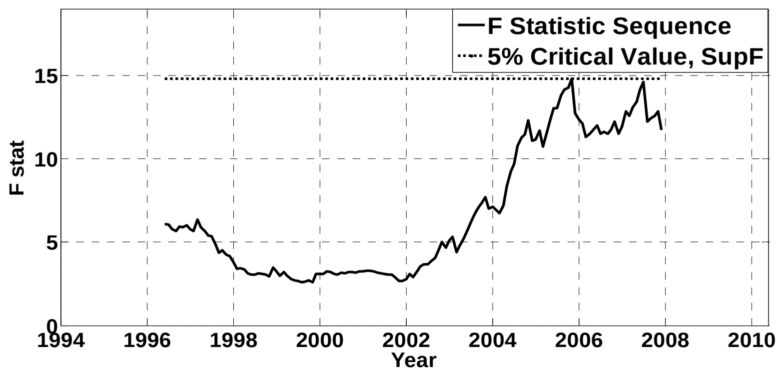

Figure 2 Constancy of the Cointegration Parameters,Japan: 1994:01-2010:04

[9] See Hansen(1992, 2002) for further detail.

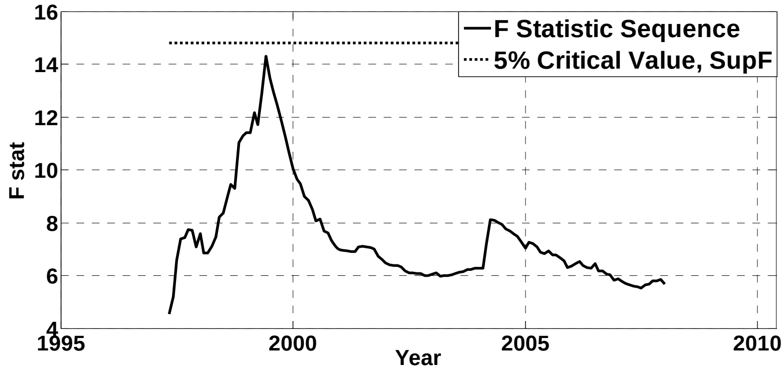

Figure 3 Constancy of the Cointegration Parameters,Korea: 1995:01-2010:04

Figures 1 to 3 outline the sequences of $F_{t/T}$ in the interval [0.15, 0.85]. The tests do not reject the null hypothesis of cointegrating parameters instability at the 5% level for all three countries, indicating that $\ln D_t$, $\ln F_t$ and $\ln P_t$ have a long-run and stable cointegrating relationship.

Based on the estimated parameters (implied IES $1/\hat{\rho}$ and $1/\hat{\upsilon}$ in Table 3), GMM is used to estimate the parameters of habit formation of import and domestic consumptions. The results are given in Table 4. In addition, the following vectors are used as instruments: constant, trend, D_t/D_{t-1} , F_t/F_{t-1} , P_{t-1}/P_t and $1+r_{t-1}$.

Following Amano and Wirjanto (1996), we set $\beta = 0.99$, and the consistent HAC covariance matrix is given by Newey- West (1987), while the weight of the auto-covariance is given by Quadratic Spectral (QS) kernel. $J-test$ is Hansen's (1982) test for overidentifying restrictions, asymptotically chi^2 distributed with n degrees of freedom, where n is the number of overidentifying restrictions and is equal to ten for all models. $Wald_{\gamma=\delta=0}$ is a test for the existence of habit formation with a null hypothesis $H_0 : \gamma = \delta = 0$. The corresponding *p*-value is included in square brackets.

Hansen's $J-test$ evaluates the extent to which the residuals are effectively orthogonal to the instrument set. It is clear that Hansen's $J-test$ does not reject the null hypothesis at the 1% level for all models, supporting the specification defined in equation (5). Simultaneously, the $Wald_{\gamma=\delta=0}$ statistics rejects the null hypothesis, $H_0:\gamma=\delta=0$, proving the significance of habit formation in most cases. This also shows the limitation of the framework introduced by Ceglowski (1991), where $\gamma=\delta=0$, and adds encouragement to our model.

The estimated parameters γ and δ from different cases are statistically significant with expected signs. The estimated coefficients are little different from each other between the two cases. In Table 4, γ is estimated to be 0.610-0.643 and δ -0.143 to -0.131 for China, indicating that imported goods for China present some durability as defined by Ferson and Constantinides (1991). Whereas γ is estimated to be 0.595-0.633 and δ 0.397-0.401 for Japan, and γ is estimated to be 0.623-0.626 and δ 0.378-0.405 for Korea. The estimated coefficients imply that in Japan and Korea, previous domestic consumption has a greater impact on current utility than previous import consumption. In addition, China has the greatest habit formation of domestic consumption among the three countries. All the parameters of habit formation would be used to estimate the modified IES in equation (10).

Table 4 Generalized method of moments (GMM) results of equation (5)

	IES values based on table 3	γ	δ	$J-test$	$Wald_{\gamma=\delta=0}$
China	$1/\hat{\rho}$=1.891, $1/\hat{\upsilon}$=1.999	0.643*** (0.012)	-0.143* (0.078)	0.118 [0.999]	3116 [0.000]
	$1/\hat{\rho}$=1.983, $1/\hat{\upsilon}$=2.132	0.610*** (0.005)	-0.131* (0.075)	0.122 [0.999]	24798 [0.000]
Japan	$1/\hat{\rho}$=0.291, $1/\hat{\upsilon}$=0.331	0.633*** (0.002)	0.401*** (0.004)	0.110 [0.999]	102169 [0.000]
	$1/\hat{\rho}$=0.252, $1/\hat{\upsilon}$=0.321	0.595*** (0.004)	0.397*** (0.002)	0.121 [0.999]	35148 [0.000]
Korea	$1/\hat{\rho}$=0.343, $1/\hat{\upsilon}$=0.635	0.626*** (0.001)	0.405*** (0.003)	0.080 [0.999]	215111 [0.000]
	$1/\hat{\rho}$=0.370, $1/\hat{\upsilon}$=0.606	0.623*** (0.002)	0.378*** (0.008)	0.067 [0.999]	137980 [0.000]

Note: Numbers in parentheses are standard errors. Numbers in square brackets stand for *p*-value. *: significant at 10%, ***: significant at 1%.

Table 5 presents the Ljung-Box and ARCH-LM tests for the residuals from GMM estimation. Ljung-Box *(p)* is for *p*th-order serial correlation in the residuals of an MA model. Ljung and Box (1978)'s modified $Q^*(m)$ statistic is introduced by Box and Pierce (1970) to increase the power of the test. Ljung-Box statistics is given by $Q^*(m) = T(T+2)\sum_{l=1}^{m}\left[\hat{\rho}_l^2/(T-l)\right] \sim \chi_\alpha^2(m)$. Simulation studies suggest that $m = \ln(T)$ provides better power performance, and m is equal to five for all tests. The decision rule is to reject H_0 of absence of serial correlation if $Q(m) > \chi_\alpha^2$. ARCH(p) LM is a standard Largrangian multiplier introduced by Engle (1982) to test whether there is *p*th-order ARCH effects in the estimated residuals.

Table 5 Misspecification tests of GMM estimation

	Equation System	Ljung-Box (2)	Ljung-Box (5)	ARCH(2) LM	ARCH(4) LM
China	Equation (1)	0.410 [0.815]	0.580 [0.989]	0.048 [0.976]	0.198 [0.995]
	Equation (2)	5.772 [0.056]	34.353 [0.000]	34.766 [0.000]	40.289 [0.000]
	Equation (1)	0.156 [0.925]	0.211 [0.999]	0.079 [0.961]	0.198 [0.995]
	Equation (2)	4.808 [0.090]	32.314 [0.000]	33.840 [0.000]	38.873 [0.000]
Japan	Equation (1)	1.841 [0.398]	0.530 [0.767]	0.039 [0.981]	0.072 [0.999]
	Equation (2)	1.978 [0.852]	4.783 [0.443]	0.579 [0.749]	1.422 [0.840]
	Equation (1)	1.820 [0.403]	4.327 [0.503]	0.164 [0.921]	0.343 [0.987]
	Equation (2)	1.076 [0.584]	3.291 [0.655]	0.452 [0.800]	0.671 [0.955]
Korea	Equation (1)	2.068 [0.356]	2.507 [0.775]	0.152 [0.927]	0.387 [0.984]
	Equation (2)	0.708 [0.871]	0.749 [0.980]	0.018 [0.991]	0.029 [0.999]
	Equation (1)	0.085 [0.959]	1.605 [0.901]	0.188 [0.910]	0.336 [0.987]
	Equation (2)	0.424 [0.809]	1.453 [0.918]	0.037 [0.982]	0.058 [0.999]

Note: Numbers in square brackets stand for *p*-values.

These test results suggest that there is no serial correlation and ARCH effects in the estimated residuals for Japan and Korea, and it is also true when we come to test the first equation for China. Whereas, there are serial correlation and ARCH effects in the second equation for China. In short, according to the two tests, serial correlation and ARCH effects do not affect our GMM estimation seriously. Therefore, the estimated results of GMM are credible and reliable.

In order to derive the relationship between import consumption and domestic consumption, we have to analyse the substitution effect between the two types of goods. Since the share of spending on domestic goods (*s*) is time-varying, the Marshallian price elasticity of imported goods calculated by equation (8) is in the range of -2.037 to -1.887 for China, -0.392 to -0.347 for Japan and -0.748 to -0.676 for Korea. The price elasticity is also time-varying with the change of *s*.

Table 6 Price and expenditure elasticities for domestic and imported goods

	Type of Goods	Average Price Elasticity	Average Expenditure Elasticity	Nature of goods
China	Imports	-1.976	1.061	Luxury
	Domestic	——	0.995	Necessity
Japan	Imports	-0.357	1.189	Luxury
	Domestic	——	0.992	Necessity
Korea	Imports	-0.708	1.571	Luxury
	Domestic	——	0.903	Necessity

As presented in Table 6, the average price elasticity of imported goods is -1.976, -0.357 and -0.708 respectively for China, Japan and Korea. The estimated average price elasticities are different from those in Kee, Nicita and Olarreaga (2008), whose estimated import demand price elasticity is -2.54 based on HS six digit and -1.12 based on ISIC three digit for China, -4.05 based on HS six digit and -1.23 based on ISIC three digit for Japan and -2.08 based on HS six digit and -1.10 based on ISIC three digit for Korea. However, all the results suggest that a decline in the relative price between imported and domestically produced goods would tend to raise the demand for imported goods in all the three countries, especially in China.

We then analyze different consumer behaviors of pursuing import and domestic goods.

By doing so, the expenditure elasticities of import and domestic goods from equations (8) and (9) are derived. As reported in Table 6, the average expenditure elasticities of imported goods are 1.061, 1.189 and 1.571 respectively for China, Japan and Korea, and the corresponding average expenditure elasticities of domestically produced goods are 0.995, 0.992 and 0.903. These results mean that imported goods are on average luxurious, but domestically produced goods are necessities.

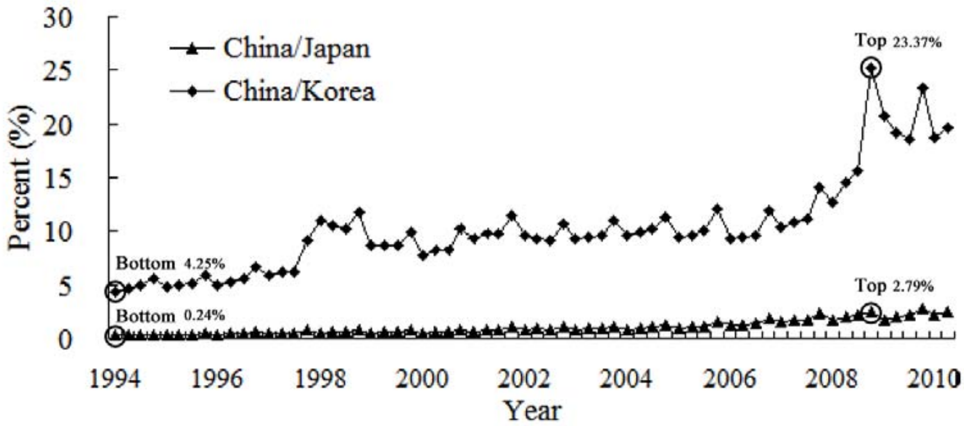

Figure 4. The Ratio of Average Per-capita Disposable Income

Next, we continue to analyze the characters of consumer behavior. China has the largest IES and modified IES of both import and domestic consumptions, partly because Chinese consumers are more vulnerable to liquidity constraints than their Japanese or Korean counterparts (Table 7). As seen in Figure 4, the average per-capita disposable income in China is significantly less than that in Korea and Japan. Therefore, the optimal intertemporal consumption pattern for Chinese consumers is easily disrupted by liquidity constraints.

Furthermore, the IES of import consumption is larger than that of domestic consumption, because domestic goods act as a necessity, while imported goods as a luxury. However, habit formation of domestic consumption is larger than import consumption, and imported goods for China even present some durability.

Table 7 Comparisons of consumer behavior in different countries

Type of consumption	China	Japan	Korea
IES of imports	2.066	0.326	0.621
IES of domestic goods	1.937	0.272	0.357
AES between import and domestic goods	0.938	0.832	0.575
Modified IES of imports	2.617	0.105	0.203
Modified IES of domestic goods	0.202	0.032	0.039
Modified AES	0.077	0.307	0.193
Habit formation of imports	-0.137	0.399	0.392
Habit formation of domestic goods	0.627	0.614	0.625
Relationship Import/domestic goods	complement	substitute	substitute

Notes: IES = intertemporal elasticity of substitution, AES = intratemporal (or intraperiod) elasticity of substitution between imports and domestic goods.

Constantinides (1990) argues that habit formation introduces a gap between IES and modified IES, and the latter is about one fourth of the size of the former. However, Naik and Moore (1996) find the gap between the two elasticities to be about one half. Moreover, Ferson and Constantinides (1991) and Ogaki and Park (1997) point out that a relatively low modified IES is compatible with a relatively high IES when habit formation is allowed. Croix and Urbain (1998) show that IES of domestic consumption is five times larger than the modified IES and IES of import consumption is nearly three times larger than the modified IES for the USA. Our estimated results prove that IES of domestic consumption is nearly nine times as large as the modified IES, while the two elasticities of import consumption are almost the same for China. But for Japan and Korea, IES of import consumption is about two times larger than the modified IES. The results reveal that habit formation plays an essential role in affecting consumer behavior.

Finally, whether import consumption crowds out domestic consumption needs to be addressed. The modified IES of domestic consumption ($1/\tilde{\rho}$) is 0.202 and the AES ($\tilde{\upsilon}/\tilde{\rho}$) is 0.077 for China (Table 7). IES is greater than AES ($1/\tilde{\rho} > \tilde{\upsilon}/\tilde{\rho}$). The results support the argument that imports and domestically produced goods are complements rather than substitutes in China. This has a critical policy implication as far as currency appreciation is concerned, as it implies that imported goods may have little crowding out effect on domestically produced goods caused by a decline in the relative price between these two types of goods.

However, due to the high IES of domestic consumption and the existence of habit formation, intertemporal consumption optimization implies that a decline in intratemporal consumption would increase the implied per-capita income, which would increase the demand for imported goods as well as domestic goods in the current period through an income effect. The bigger IES, the more will be consumed in the current period at the expense of future consumption. This income effect is opposite to the intratemporal substitution effect. Since IES is bigger than AES, there is no crowding out effect on current domestic demand.

In addition, imported goods present some durability and substitute little for domestic goods in China. That is also why the modified AES is only 0.077 as compared to 0.307 for Japan and 0.193 for Korea. Thus domestic consumption is little influenced by intratemporal optimality choice when the relative price of imports and domestically-produced goods declines. This is why imports and domestic consumptions act as complements to each other.

In contrast, IES is smaller than AES in Japan and Korea, implying that import and domestic consumptions are substitutes in both countries. This may be explained as follows. Firstly, as Japan and Korea have a good medical and insurance system, unlike their Chinese counterparts, Japanese and Korean consumers are less vulnerable to liquidity constraints. As a result, IES and modified IES in Japan and Korea are smaller than those in China. Therefore, the intertemporal substitution effect in Japan and Korea is not as strong as in China. Secondly, Table 7 shows that the average expenditure elasticities of import consumption in Japan and Korea are greater than that in China, indicating that Japanese and Korean consumers would spend more on imported goods than their Chinese counterparts as a result of rising per capita incomes. Thirdly, the ratio of average per-capita disposable incomes between China and Japan was only about 2.8% and that between China and Korea 20% in 2009 (Figure 4). This means that a decline in the relative price between imported and domestically produced goods would sharply raise import consumption in Japan and Korea due to an income effect, strongly crowding out domestic consumption because of an intratemporal optimality choice. That is also why the modified AES in Japan and Korea are much greater than in China.

6. Conclusion

In this paper, we employ a two-good permanent-income model to investigate whether imports crowd out domestic consumption in China, Japan and Korea.

We take full advantage of the well-developed theory of cointegration to investigate IES of both import and domestic consumptions, pursue GMM approach to estimate the habit formation parameters, and calculate the modified IES and modified AES on habit formation.

The modified IES of domestic consumption are estimated to be 0.202, 0.032 and 0.039 for China, Japan and Korea, respectively, and the corresponding modified IES of import consumption 2.617, 0.105 and 0.203. The estimated AES are 0.077, 0.307 and 0.193 respectively for China, Japan and Korea.

As the IES between import and domestic consumptions is greater than the AES in China, it suggests that import and domestic consumptions are complements. In Japan and Korea, the IES is smaller than the AES, suggesting that import and domestic consumptions are substitutes. These results imply that the crowding out effect of imports on domestic consumption is limited in China but strong in Japan and Korea.

Three possible explanations are offered for the different results between China and the other two countries. First, China's per capita income is significantly lower than that in Japan or Korea. This implies that Chinese consumers must have been more vulnerable to liquidity constraints than their Japanese or Korean counterparts. Therefore, the Chinese pay more attention to current consumption than their Japanese and Korean counterparts. Consequently, a decline in the relative price between imported and domestically produced goods leads to a rise in implied per-capita income, which would increase the demand for imported goods as well as domestic goods in the current period through an income effect. However, as AES is very small in China, the substitution effect of imports on domestic consumption is critically diluted by the income effect.

Second, since the average expenditure elasticities of import consumption in Japan and Korea are greater than that in China, compared to their Chinese counterparts, Japanese and Korean consumers tend to spend more on imported goods as a result of rising per capita disposable incomes.

Third, China has the highest IES of domestic consumption among the three countries. Compared to their Japanese and Korean counterparts, Chinese consumers tend to consume more domestically produced goods in the current period relative to such future consumption. In addition, imported goods present some durability, which makes the modified AES as small as 0.077, compared to 0.307 in Japan and 0.193 in Korea. Thus domestic consumption is little impacted by intratemporal optimality choice when the relative price between imported and domestically produced goods declines.

Our results have striking policy implications for China relating to currency appreciation. As habit formation is an important element in consumer behavior, it reduces IES in a big scale. This suggests that the modified IES is important for investigating consumer behavior of intertemporal substitution choice. It also reveals the limitations in the framework introduced by Ceglowski (1991), Clarida (1994), Amano and Wirjanto (1996) and Xu (2002), where all parameters of habit formation are set to zero.

Compared with China, domestic consumption in Japan and Korea is more sensitive to the relative price between imported and domestically produced goods. In addition, our empirical results imply that import and domestic consumptions are complements for China. Therefore, China should continue to speed up the pace of opening-up and develop international trade. However, one should not be over optimistic, as the consumption capability of Chinese consumers would depend on a steady increase of their average disposable incomes. If import consumption contained less luxurious goods compared to domestic consumption, there would be no difference between imported and domestically produced goods in China. Consequently, intratemporal substitution effects would increase, reducing the degree of complementarities between import and domestic consumptions.

Appreciation of the Chinese currency would have this anticipated effect as it will reduce the relative price between imported and domestically produced goods. In the short run, the crowding out effect of imports on domestically produced goods may be limited due to a low intratemporal substitution effect. In the long term, however, the situation may change, especially when per capita income in China rises. In that case, China's consumption habit may approach that of Japan's or Korea's, meaning that the crowding effect of imports on domestically produced goods will increase over time.

References

Amano, R. A., and Wirjanto, T. S., 1996. Intertemporal substitution, imports and permanent income model. Journal of International Economics 40, 439-457.

Amano, R. A., Ho, W. M., and Wirjanto, T. S., 1998. Intraperiod and intertemporal substitution in import demand. Cahiers de recherche CREFE Working Paper.

Boldrin, M., Christiano, L., and Fisher, J., 1995. Asset pricing lessons for modeling business cycles. NBER Working Paper NO. 5262.

Box, G. E. P., and Pierce, D. A., 1970. Distribution of residual autocorrelations in autoregressive-integrated moving average time series models. Journal of the American Statistical Association 65, 1509-1526.

Ceglowski, J., 1991. Inter-temporal substitution in import demand. Journal of International Money and Finance 10, 118-130.

Clarida, R. H., 1994. Co-integration, aggregate consumption, and the demand for imports: a structural econometric investigation. The American Economic Review 84, 298-308.

Constantinides, G. M., 1990. Habit formation: a resolution of the equity premium puzzle. Journal of Political Economy 98, 519-543.

Cooley, T. F., and Ogaki, M., 1991. A time series analysis of real wages, consumption, and asset returns under optimal labor contracting: a cointegration-Euler equation approach. Rochester Center for Economic Research Working Paper No. 285, Rochester, New York: University of Rochester.

Croix, D. L., and Urbain, J. P., 1998. Inter-temporal Substitution in Import Demand and Habit Formation. Journal of Applied Econometrics 13, 589-612.

Duesenberry, J. S., 1949. Income, Saving and the Theory of Consumer Behavior. Harvard University Press, Cambridge, MA, USA.

Eichenbaum, M. S., Hansen, L. P., and Singleton, K. L., 1988. A time series analysis of representative agent models of consumption and leisure choice under uncertainty. Quarterly Journal of Economics 103, 51-78.

Engle, R., 1982. Autoregressive conditional heteroskedasticity with estimates of the cariance of U. K. inflation. Econmetrica 50, 987-836.

Engle, R., and Granger, C., 1987. Cointegration and error correction: representation, estimation and testing. Econometrica 55, 251-276.

Ferson, W. E., and Constantinides, G. M., 1991. Habit persistence and durability in aggregate consumption. Journal of Financial Economics 29, 199-240.

Hansen, B. E., 1992. Tests for parameter instability in regressions with I(1) processes. Journal of Business and Economic Statistics 10, 321-335. Reprinted in Twentieth Anniversary Commemorative Issue. Journal of Business and Economic Statistics (2002) 20, 45-59.

Hansen, L., 1982. Large sample properties of generalized method of moments estimators. Econometrica 50, 1029-1054.

Kee, H. L., Nicita, A., and Olarreaga, M., 2008. Import demand elasticities and trade distortions. The Review of Economics and Statistics 90, 666-682.

Ljung, G. M., and Box, G. E. P., 1978. On a measure of lack of fit in time series models. Biometrika 65, 297-303.

MacKinnon, J. G., 1996. Numerical distribution functions for unit root and cointegration tests. Journal of Applied Econometrics 11, 601-618.

Muellbauer, J., 1988. Habits, rationality and myopia in the life cycle consumption function. Annales d'Economie et de Statistique 9, 47-70.

Naik, N.Y., and Moore, M. J., 1996. Habit formation and intertemporal substitution in individual food consumption. The Review of Economics and Statistics 78, 321-328.

Newey, W., and West, K., 1994. Automatic lag selection in covariance matrix estimation. Review of Economic Studies 61, 631-653.

Nieh, C. C., and Ho, T. W., 2006. Does the expansionary government spending crowd out the private consumption? cointegration analysis in panel data. Quarterly Review of Economics and Finance 46, 133-148.

Nishiyama, S. I., 2005. The cross-Euler equation approach to inter-temporal substitution in import demand. Journal of Applied Econometrics 20, 841-872.

Ogaki, M., 1992. Engel's law and cointegration, Journal of Political Economy 100, 1027-1046.

Ogaki, M., and Park, J. Y., 1997. A cointegrating approach to estimating preference parameters. Journal of Econometrics 82, 107-134.

Phillips, P., and Hansen, B. E., 1990. Statistics inference in instrumental variables with I(1) processes. Review of Economic Studies 57, 99-124.

Stock, J. H., and Watson, M. W., 1993. A simple estimator of co-integrating vectors in higher order integrated systems. Econometrica 61, 783-820.

Xu, X., 2002. The dynamic-optimization approach to import demand: a structural model. Economic Letters 74, 265-270.