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THE ASYMMETRIC EFFECT OF INCOME ON IMPORT DEMAND IN GREECE.

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ABSTRACT

This paper presents empirical evidence supporting the argument that a significant asymmetry exists in the income elasticity of Greek imports. Using multivariate cointegration techniques for the estimation of long-run imports we derive short-run error correction equations that separate income elasticities for periods when income is rising and periods when it is falling. The empirical results show that the response of imports to rising income is stronger than the response of imports to falling income. The important policy implication of this asymmetry is that a consecutively positive growth would lead imports to continuously increase causing the current account deficit to persistently widen.

JEL classification: F14, C22, C32, C51

Keywords: Import demand, asymmetric income elasticities, multivariate cointegration, short-run error correction, current account deficit, structural reforms.

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1. Introduction

The recent improvement in Greece's current account deficit took place after a prolonged recession with a sharp correction being witnessed after five years of recession. One of the reasons the deficit is likely to improve during a recession is that a decrease in consumer and investment spending implies a decline in imports to the extent that the goods and services purchased are not domestically produced.

This paper offers an empirical estimation of import demand, attempting to explain the above-mentioned slow adjustment of the external sector, providing evidence that could provide some pointer to its future development. Aiming at contributing to the literature on import demand, we show that there exists a statistically significant asymmetry in the income elasticity of imports that explains imports' slow adjustment during the recent recession. The estimation method we adopt allows us to estimate income elasticities for periods when income is rising and periods when it is falling. If the response of imports to rising income is stronger than the response of imports to falling income, then the net effect on imports will be positive.

The important policy implication of this asymmetry is that a continuous period of positive growth would lead imports to continuously increase causing the current account deficit to continuously widen. This widening would eventually necessitate a severe recession in order to reestablish equilibrium and reduce the value of external debt. Furthermore, *ceteris paribus*, the beneficial effect of a contraction on the balance of payments deficit would be quickly offset by an expansion.

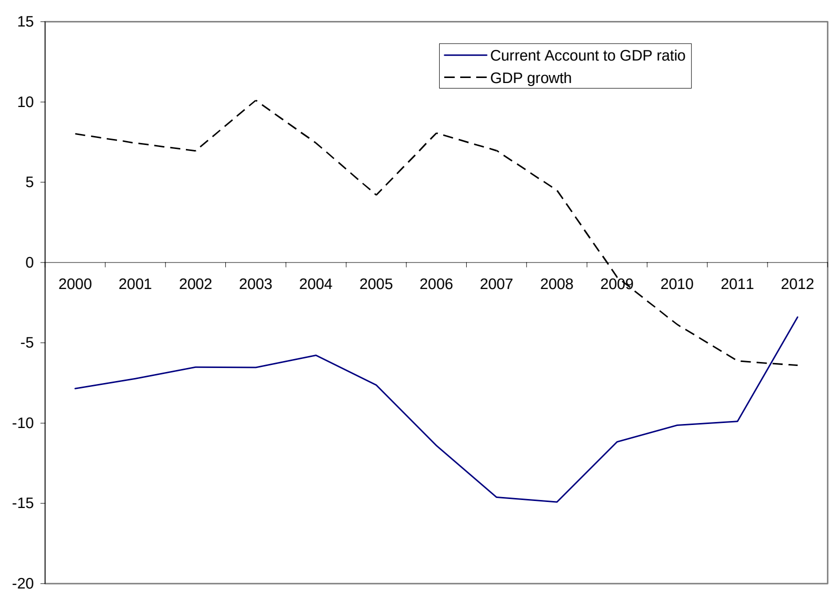
We adopt the Johansen procedure to estimate the long-run import demand function and the general-to-specific methodology to derive the short-run determinants of import demand. We test for asymmetric income effects using quarterly data covering the past eleven years. The assumption of asymmetric income effects in the short run as opposed to symmetry constitutes an improvement based on a number of tests of robustness. Income elasticities with respect to expansions are found to be well above unity while income elasticities with respect to contractions are below unity. Moreover, according to this evidence it can be seen that failure to assume asymmetry explains why some researchers have presented low income elasticities in the short-run (see for example Athanasoglou 2010).[1]

The outline of the paper is as follows. Section two provides some descriptive statistics on the balance of payments, a brief survey of the literature and presents the empirical specification. In section three, the estimation results are presented. Finally, section four provides conclusions and policy recommendations.

2. Background issues and the estimation model.

As noted in the introduction, the Greek current account deficit has been improving recently but at a rather slow pace. In 2008, Greece reported a current account deficit of 14.9 % of the country's GDP which is considered to be rather high when historically this measure averaged 5.5%. In 2010 and 2011, the current account deficit as a percentage of GDP showed little improvement at

Figure1. Current Account to GDP ratio and GDP growth



[1] Income elasticities in the long-run are usually found to be high close or above unity (see for example Hooper, Johnson and Marquez 2000, Table 1 in p.8).

10.1% and 9.9% respectively, even though the country had been in a recession since 2008. The figure showed a more significant improvement in 2012, at 3.4%. This development can be attributed mainly to the PSI effect (which reduced net interest payments on Greek government bonds held by non-residents) and the reduction in imports which account for more than three times the value of the country's exports (if one takes the average for 2000-2010).[2] Assessing the response of the import bill to a future expansion in the economy is therefore critical for determining current account sustainability.

The empirical determinants of import demand found in early work have been extensively reviewed by Goldstein and Khan (1985).These earlier studies focused on either absolute or relative prices and the estimation method was largely standard ordinary least squares (OLS). More recent studies use more sophisticated techniques to perform similar estimates. Zonzilos (1991) uses the Engle-Granger method to estimate long- and short-run import demand in Greece adopting absolute prices, while Milas (1998) and Athanasoglou (2010) use the Johansen procedure and find a long-run import demand for Greece adopting the relative price specification. Greek import demand elasticities are estimated as well by Sinha and Sinha (2000) who use the Phillips-Hansen fully modified estimate. Chang, Ho and Huang (2005) use bounds test analysis by Pesaran to estimate a relative price version of import demand for Korea. Long-run income elasticities are in most studies high, above or close to unity but short-run elasticities where they are estimated are found to be rather low. This may be due to not allowing for asymmetric effects.

In the trade literature, the issue of a potentially asymmetric response of imports to GDP has generally received little attention. Most of the existing empirical research, which is scarce, examines the response of prices to variables such as cost or exchange rates (see Zombanakis 1998 p.1). Asymmetry is more often encountered in the estimation of energy demand (with respect to prices and income, see, for example, Gately and Huntington 2001) and few studies exist in the consumption literature (Till van Treeck 2008).

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[2] To look at the issue in a different light, they represented about 75% of total trade during this period.

The most common specification of the import demand function uses a measure of income (GDP or industrial production), import prices and domestic prices as independent variables. In our specification, we test the hypothesis of homogeneity with respect to prices[3] to see whether using the price ratio constitutes an improvement. The hypothesis fails to be rejected with a χ^2 of 0.079 (p-value: 0.779). The relative price specification is therefore adopted and is augmented with capacity utilization. The two equations are as follows:[4]

$$m_t = f\ (ip,\ ln(p_{m,t}/p_{d,t})) \tag{1}$$

$$m_t = f\ (ip,\ ln(p_{m,t}/p_{d,t}),\ cu_t) \tag{2}$$

where m_t are real imports, ip_t industrial production, p_m and p_d import and domestic prices, respectively, and cu_t capacity utilization. Capacity utilization is often introduced into the import demand equation in order to capture the potential for import substitutability by taking cyclical effects into account. Through its inclusion, the equation is more corrected specified and omitted variable bias is avoided, thus permitting more accurate estimates of the elasticities. It has been shown that not allowing for cyclical effects leads to rather high (i.e., greater than one) income elasticities in the long-run, while their inclusion produces unitary income elasticities (see Athukorala and Menon, 1995).

3. Data and empirical results

Real imports are non-oil imports in millions of euro divided by the import price index. Import prices and domestic prices (producer prices) are both indices, using 2005 as a base year. Industrial production and capacity utilization are also indices. The data is quarterly, refers to the period 2000 to 2011 and is produced by the Bank of Greece except for industrial production which is obtained by ELSTAT.

[3] The test was applied to the long-run equation estimating with the Johansen procedure and with one cointegrating vector.

[4] Lower case letters denote variables expressed in logarithms.

We first test for the stationarity of the variables and estimate the long-run model. The results of the ADF testing for the presence of a unit root are presented in Table 1. The p-values show that all series are I(1). We then proceed to estimate the two versions of the import demand equations (1) and (2) using the Johansen procedure. Table 2 reports the results. The top half of the table presents the estimated eigenvalues and the trace and maximum eigenvalue statistics along with the critical values of the tests of hypotheses on the value of *r*, the number of cointegrating vectors. A * indicates rejection of the null shown on the left hand-side of the table at the 10% level of significance. The procedure identifies one cointegrating vector according to the trace test for equation (1) and the maximum eigenvalue test for equation (2).[5] In addition, we look at the roots of the companion matrix to strengthen the above result. For both equations, all the roots but the first lie inside the unit circle and the second root is substantially smaller than the first root.[6] All the above evidence indicates the existence of one cointegrating vector in both cases.

The lower half of Table 2 shows the estimated cointegrating vectors along with the corresponding weights α_i and a test statistic of the weak exogeneity of relative prices. The hypothesis of weak exogeneity of relative prices fails to be rejected at 10% level of significance and, thus, the restriction is imposed in the estimation. We observe that the first vector contains the assumed import demand relation with the correct signs and all coefficients are significant as the relevant t-statistics show at 5% level of significance. The weights are highest in the first equation of the system but also large in the second equation, indicating the endogeneity of industrial production which is captured by the procedure improving the efficiency of the estimates producing a smaller mean square error.

In the second stage, where the dynamic short-run error correction equation for imports is estimated, the asymmetric response to income is incorporated. These equations represent a reparameterization according to the general-to-specific method, where

[5] Further support for the presence of one cointegrating vector is offered by the existence of a large difference in size between the first and the second eigenvalue.

[6] Results are available upon request.

insignificant variables are discarded and a more parsimonious specification adopted. First, the short-run error correction equation assuming symmetry is estimated: [7]

$$\Delta m_t = -0.006 + 0.567\Delta ip_t - 1.660\Delta ln(p_m/p_d)_t - 0.190EC_{t-1} +$$

$$(-0.908) \quad (3.017) \quad (-2.223) \qquad\qquad (-1.563)$$

$$0.083{*}DGP - 0.124\ DC1 + -0.001\ DC2{*}\ trend$$

$$(4.456) \qquad (-3.791) \qquad (-2.206) \qquad\qquad\qquad (1.1)$$

$$R^2=0.683 \quad \sigma=0.030 \quad LM4=0.855[0.601] \quad F(RESET)=0.948[0.336]$$

$$\chi^2_{2,JB}=1.621[0.444] \quad F(ARCH(4))=0.613[0.655]$$

R^2 is the coefficient of determination and σ the estimated standard error,

LM4 is the Lagrange multiplier F-test for serial correlation of up to order 4,

F(RESET) is Ramsey's RESET test for correct specification of the equation,

$\chi^2_{2,JB}$ is the Jarque-Bera test for normality of the residuals which is distributed with Chi-square distribution with 2 degrees of freedom,

Finally, F(ARCH(4)) is the fourth order Lagrange multiplier test of autoregressive conditional heteroscedasticity.

DC1 and DC2 are dummy variables capturing the recent economic crisis. It is expected that the estimated coefficients will not be constant over the sample especially during the years of the economic crisis.

This simple specification shows that if we wrongly assume that demand is perfectly income "reversible" (i.e., that the effect of increases and decreases is symmetric), income elasticity is rather low, while the relative price elasticity is higher than what is usually found in previous work on import demand estimation (see for example, Zonzilos, 1991, Athanasoglou, 2010 and Hooper, Johnson and Marquez, 2000, Table 2, p.9).

[7] DGP is a dummy variable that refers to government policy regarding increases in prices of domestically produced goods during certain periods in response to the rising cost of major raw materials.

We then introduce asymmetry by decomposing income changes into increases denoted by Δipp_t and decreases denoted by Δipn_t (to positive and negative). Estimation yields equations (1.1a) and (2.1) in Table 3. None of the diagnostic tests for these equations is statistically significant - the residuals follow an innovation process free from autocorrelation and conditional heteroscedasticity. For equation (2.1) the ARCH statistic for conditional heteroscedasticity is significant at 1% but not significant at 0.5%.

They are also normally distributed and they indicate the correct specification of both the long-run and the short-run equations (including the weak exogeneity hypothesis). A comparison with the diagnostics of equation (1.1) shows that the above two versions constitute an improvement.

An F-test of equality of the coefficients of positive and negative income changes of 11.69 and 19.73 for (1.1a) and (2.1), respectively, leads to the rejection of the hypothesis of a symmetric response from imports. Instead, imports are more responsive to income increases than to income decreases or the response of imports to positive changes in income is larger than that of negative changes. The asymmetry that arises always leaves a positive net effect on imports over the cycle and leads to continuous increases in imports and a widening of the the trade deficit assuming exports remain unchanged. In versions (1.1a) and (2.1), the coefficients of the positive income changes are very close indicating that a 10% increase in income causes a 15% increase of imports accompanies, while the corresponding effect when income decreases by 10% is between 5% and 7% (one third or half approximately of the response to increases). A consequence of wrongly assuming symmetry regarding efficiency of the estimated elasticities can be seen by comparing (1.1) and (1.1a) and (2.1). In the model where import demand is perfectly "income reversible"[8], the estimated income elasticity is biased downwards while the relative price elasticity is overestimated.

These results have important policy implications: They are consistent with the slow adjustment of the trade balance to recessions and suggest that the negative effect on the trade balance of an expansion is reversed only after a severe recession.

[8] A similar result of downward bias was produced by Gately and Huntington (2001) regarding the response of energy demand to income when symmetry is wrongly assumed.

To further investigate the sensitivity of import demand to income changes, we consider differences in response depending on the magnitude of income growth. That is, does it matter if we are faced with a relatively larger or a smaller expansion. This implies a further decomposition of Δipp_t to increases that are large and refer to a bigger expansion and to smaller increases that refer to lower growth rates $\Delta ipph_t$ and $\Delta ippl_t$ respectively. The cut-off point that classifies an increase to one or the other group is the mean rate of increase. Estimation with this further breakdown yields equations (1.2) and (2.2) in Table 3.

By applying this extension, another aspect of the analysis regarding asymmetric effects emerges. Asymmetry is also related to the size of the increase in the rate of growth. Income elasticities are larger when growth rates are low leading to the conclusion that even a small expansion can trigger a significant import increase and a deterioration of the trade balance. At higher growth rates, income elasticities are still high (higher than those associated with income decreases).

In a final modification of the short-run specification, we consider another breakdown of asymmetric income effects related to the timing of an expansion; that is whether it follows a recession or a previous expansion. The income growth variable is decomposed to $\Delta ipprec_t$ which represents the positive growth rates that follow negative ones and to $\Delta ippexp_t$ the positive growth rates that follow expansions. The results are presented in equations (1.3) and (2.3) in Table 3.

Income elasticities are larger when we consider increases in income that come after a period of expansion compared to increases that follow a contraction. This result can also be verified by observing that the values of growth rates that correspond to consecutive growth in most cases are small in magnitude. According to our previous finding, these rates will tend to produce a larger response. It then follows that the trade balance will deteriorate more intensely in periods of consecutive growth.

4. Conclusions

The empirical estimation of import demand focusing on asymmetric income effects has revealed the following:

- The demand for imports responds more to an expansion than to a contraction of income in the short-run. Introducing asymmetry in the income effect leads to an elasticity of for an income expansion of 1.32, while the elasticity for an income contraction is 0.47; in both cases, the impact of capacity utilization is taken into account. If this variable is not included in the analysis then the corresponding expansion and contraction elasticities are 1.46 and 0.69 respectively.
- The estimated short-run elasticities of import demand with respect to relative prices when we separate the effect of an expansion from that of a contraction are close to unity (-1.094 if capacity utilization is included and -1.161 if this variable is not included). In the relevant literature, estimates of this variable have been found to be close to unity or inelastic.[9]

The empirical analysis was designed to illustrate the existing asymmetry in the response of imports to income changes. Providing robust estimation results, we have shown that the negative effect on imports that is produced by a recession is weaker than the corresponding positive effect of an expansion. This is consistent with the hypothesis that the impact of the recent recession on the trade balance has been rather slow due to the delay in the implementation of structural reforms. The low level of foreign direct investment and the failure to substitute imports represent a continuous threat to the sustainability of the current account improvement. The consequence of policies aimed at attracting FDI and enhancing import substitution will be an improvement in the financing of the current account, and will increase employment.

[9] For example, close to unity absolute price elasticities are estimated by Zonzilos (1998) and Sinha and Sinha (2000). Chang, Ho and Huang (2005), Milas (1998) and Athanasoglou (2010) find import demand to be inelastic with respect to relative prices.

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Table 1: p-Values of ADF Unit Root Tests

Variables	Levels			First Differences		
	No intercept or trend	Intercept	Trend and intercept	No intercept or trend	Intercept	Trend and intercept
m_t	0.122	0.996	0.991	0.001	0.606	0.000
ip_t	0.326	0.991	0.999	0.459	0.000	0.000
$ln(p_{mt}/p_{dt})$	0.151	0.446	0.888	0.001	0.000	0.202
cu_t	0.161	0.952	0.876	0.000	0.000	0.000

Note: All data are seasonally adjusted with the ratio to moving average method.

Table 2: Cointegration analysis of import demand 2000:1 2011:4

Maximum likelihood tests										
Equations / Variables		(1)				(2)				
		λ_i: 0.352 0.194				λ_i: 0.527 0.265 0.124				
H_0:	H_1:	*trace*	*trace (0.90)*	λ_{max}	λ_{max} *(0.90)*	*trace*	*trace (0.90)*	λ_{max}	λ_{max} *(0.90)*	
r=0	r=1	23.35*	22.76	15.60	16.74	38.03	39.12	23.93*	22.98	
r≤1	r=2	7.75	10.50	7.75	10.50	14.10	22.76	9.84	16.74	
r≤2	r=3					4.25	10.50	4.25	10.50	

Estimated cointegrating vectors		
$\boldsymbol{m_t}$	1	1
Constant	2.826 (3.338)	5.448 (9.363)
$\boldsymbol{ip_t}$	-2.565 (-13.934)	-1.521 (-12.467)
$\boldsymbol{ln(p_{mt}/p_{dt})}$	0.378 (2.040)	0.910 (12.120)
$\boldsymbol{cu_t}$	-	-1.722 (-7.580)
Estimated weights		
$\boldsymbol{\alpha_1}$	-0.678(-4.124)	-1.515 (-4.565)
$\boldsymbol{\alpha_2}$	-0.291(-3.264)	-0.135 (0.614)
$\boldsymbol{\alpha_3}$	-	-0.084 (0.829)
LR test for weak exogeneity of price ratio		
χ_1^2	4.702[0.030]	0.154[0.695]

Note: The trace and the maximal eigenvalue statistics are adjusted for degrees of freedom. Numbers in parentheses are t statistics and in brackets the probability values. All the variables are seasonally adjusted with the ratio to moving average method.

Table 3: Estimation of the short-run error correction equations of asymmetric imports
2000:1-2011:4

Equations / Variables	**1.1a**	**2.1**	**1.2**	**2.2**	**1.3**	**2.3**
Constant	-0.008 (-1052)	-0.009 (-1.507)	-0.010 (-1.283)	-0.011 (-1.790)	-0.009 (-1.639)	-0.009 (-1.479)
Δipp_t	1.461 (4.101)	1.319 (4.859)		-	-	-
Δipn_{g-2}	0.691 (1.898)	0.469 (1.603)	0.686 (1.880)	0.502 (1.765)	0.779 (2.733)	0.682 (2.351)
$\Delta ipph_t$	-	-	1.373 (3.702)	1.169 (4.227)	-	-
$\Delta ippl_t$	-	-	2.215 (2.412)	2.423 (3.573)	-	-
$\Delta ipprec_t$	-	-		-	1.244 (4.350)	1.243 (4.402)
$\Delta ippexp_t$	-	-		-	1.499 (1.215)	2.005 (1.577)
$\Delta ln(p_m /p_d)_{t-1}$	-	-1.094 (-1.832)	-	-0.880 (-1.486)	-	-1.580 (-2.649)
$\Delta ln(p_m /p_d)_{t-6}$	-1.161 (-1.718)	-	-1.181 (-1.741)	-	-1.860 (-3.277)	-
cu_t	-	0.667 (2.489)	-	0.606 (2.238)	-	0.415 (1.376)
EC_{t-1}	-0.192 (-1.588)	-0.465 (-4.023)	-0.180 (-1.482)	-0.463 (-4.236)	-0.328 (-2.826)	-0.398 (-3.175)
DGP	0.072 (4.023)	0.065 (6.026)	0.069 (3.800)	0.064 (6.018)	0.073 (5.711)	0.072 (5.652)
DC1	-0.163 (-5.257)	-0.130 (-5.054)	-0.162 (-5.236)	-0.134 (-5.336)	-0.144 (-5.704)	-0.134 (-5.153)
DC2**trend*	-0.001 (-2.118)	-0.001 (-2.988)	-0.001 (-1.915)	-0.001 (-2.847)	-0.001 (-2.514)	-0.001 (-2.642)
Diagnostics						
R^2	0.718	0.824	0.716	0.833	0.813	0.817
SER	0.029	0.022	0.029	0.022	0.023	0.023
Jarque-Bera $\chi^2(2)$	0.051[0.974]	0.274[0.872]	0.503[0.777]	0.183[0.913]	1.215[0.544]	0.486[0.784]
F(ARCH(4))	0.451[0.770]	5.071[0.003]	0.490[0.743]	2.814[0.041]	1.242[0.312]	1.456[0.237]
F(RESET)	0.965[0.333]	0.981[0.329]	0.870[0.358]	0.231[0.634]	0.715[0.404]	0.559[0.460]
LM(4)	0.728[0.570]	1.260[0.307]	0.583[0.691]	1.027[0.409]	0.149[0.962]	0.039[0.996]

Note: Δ denotes first differences, t statistics are in parentheses and p values in square brackets. SER is the standard error of the regression; Jarque-Bera is the chi-square normality test of residuals, F(ARCH(4)) is the F test for 4th order autoregressive conditional heteroscedasticity, F(RESET) is the F test for first order Ramsey's test for specification error, F(HET) is White's test for heteroscedasticity and LM(4) is the LaGrange Multiplier F test for 4th order serial correlation.

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