E C O N O M I C S B U L L E T I N

Evidence of the Long-Run Neutrality of Money: The Case of South Korea and Taiwan

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Abstract

This paper investigates the long-run neutrality of money using quarterly data of South Korea and Taiwan and the methodology of King and Watson (1997); particular attention is given to the integration and cointegration properties of the variables. Empirical evidence provides considerable support for the long-run neutrality of money with respect to real output in the case of South Korea, indicating that it is consistent with two of the business cycle, i.e, the monetary intertemporal model and the monetary misperceptions theory. There is little evidence that the long-run monetary neutrality hypothesis holds in the case of Taiwan. Based on the estimated results, the hypothesis of the short-run neutrality of money is rejected for both countries.

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Keywords: Money, Real output, Neutrality, Long-run relationship

JEL classification: C32, E51

1 Introduction

Traditional, commonly-accepted theories stress the long-run effect of a permanent change in money stock on real economic variables, a notion which has been referred to as the "monetary neutrality hypothesis". For example, the Quantity Theory of Money postulates that a permanent unit, or a stochastic shock to the money supply, has a unit, or a proportional effect on prices, but a zero effect on real output in the long-run. Over most of the past three decades, an abundance of tests on the neutrality proposition have been reported in the literature. The first method examines neutrality from a cross-country perspective using data averaged over long periods for a cross-section of countries (e.g. Lothian, 1985; Duck, 1993; Loef, 1993) The second is based on frequency-domain time series techniques (e.g. Lucas, 1980; Geweke, 1982, 1986; Mills, 1982). The third reported in Stock and Watson (1988), King and Watson (1997) and Fisher and Seater (1993), is based on explicit tests of coefficient restrictions in bivariate VARs. In these, long-run neutrality implies a zero restriction on the sum of the coefficients of the contemporaneous and lagged monetary variables in a regression on real economic activity. Yet, these studies have cast doubt on the empirical findings of earlier studies which overlook the time series properties of the underlying variables.

In fact, Lucas (1972a) and Sargent (1971) are credited with being the first to show that it is impossible to test long-run neutrality using reduced form econometric methods.[1] Taking this a step further, Fisher and Seater (1993) and King and Watson (1997) later argued that only if both nominal and real variables satisfy certain nonstationary conditions, can meaningful neutrality tests be constructed. The reason for this is that monetary neutrality involves a permanent change in the level of money stock, and we cannot effectively test such a theory without solid evidence that the actual money stock has been affected by a permanent change. In this sense, nominal and real variables need to be integrated of order one, but not cointegrated (Engle and Granger, 1987). If the real output and money series are nonstationary and cointegrated, then a finite vector autoregressive (VAR) process in first difference does not exist, and this typically means that long-run neutrality must be flatly rejected.

Fisher and Seater (1993, hereafter FS) defined a certain long-run derivative (LRD) for testing the long-run neutrality proposition. King and Watson (1997, hereafter KW) proposed an eclectic approach and attested to its robustness in testing neutrality. Since their contributions in this regard, many studies have adopted their methods to test the neutrality hypothesis. To cite some examples, Boschen and Otrok (1994), Serletis and Krause (1996), Olekalns (1996) and Coe and Nason (2003, 2004) have adopted FS's approach to test long-run monetary neutrality for certain

[1]See also McCallum (1984) and King and Watson (1997) for detailed explanations.

OECD countries, namely the U.S., Australia, Canada and Germany. Bullard and Keating (1995) employed the Blanchard and Quah (1989, hereafter BQ) approach to investigate the long-run relationships between inflation and output. Koustas (1998) and Serletis and Koustas (1998) adopted KW's method and, in so doing, found a groundswell of international evidence in support of the neutrality of money. See Bullard (1999) for a thorough review on this issue.

The aim of this paper is to test the long-run neutrality of money for South Korea and Taiwan by employing the improved econometric methodology of King and Watson (1997). They obtained their measures of the long-run real output responses to a permanent monetary shock in the context of a bivariate VAR framework, where only a minimal amount of structure is assumed, in fact, merely enough to identify the structural shocks of interest. Here, we implement the robustness checks put forth by King and Watson (1997) and calculate the measures of long-run real output responses for a wide range of assumed identifying parameter values. This allows us to clearly see how identifying assumptions can be mapped into estimates of long-run real output responses to a permanent monetary shock. The investigation in this paper is interesting because most of the empirical studies of the neutrality of money have tested the hypothesis for one or a group of industrialized countries, with little attention given to developing countries.

The organization of this paper is as follows. Section 2 discusses the econometric methodology. Section 3 presents the data and the estimated results. We first show the integration and cointegration properties of money and real output, and then we move to our results based on our three identifying restrictions. In Section 4, we summarize the conclusions that we draw from this research.

2 Methodology

Let m_t and y_t be the natural logarithms of the money supply and real output at time t, respectively, written in the first difference form so that both m_t and y_t are integrated of order one, but not cointegrated. In step with King and Waston (1997), we consider the following bivariate VAR(p) model:

$$\Delta m_t = \lambda_{my}\Delta y_t + \sum_{j=1}^{p} \alpha_{j,my}\Delta y_{t-j} + \sum_{j=1}^{p} \alpha_{j,mm}\Delta m_{t-j} + \varepsilon_t^m, \tag{1}$$

$$\Delta y_t = \lambda_{ym}\Delta m_t + \sum_{j=1}^{p} \alpha_{j,yy}\Delta y_{t-j} + \sum_{j=1}^{p} \alpha_{,ym}\Delta m_{t-j} + \varepsilon_t^y, \tag{2}$$

where ε_t^m and ε_t^y are, respectively, the monetary and real output structural shocks that can have permanent effect on the level of the endogenous variables, m_t and y_t. We do not provide the constant terms for expositional convenience.

We can rewrite the system in a more compact way as:

$$\alpha(L)\boldsymbol{X}_t = \boldsymbol{\varepsilon}_t, \tag{3}$$

where

$$\alpha(L) = \sum_{j=0}^{p} \alpha_j L^j$$

and

$$\boldsymbol{X}_t = \begin{bmatrix} \Delta m_t \\ \Delta y_t \end{bmatrix}, \quad \boldsymbol{\varepsilon}_t = \begin{bmatrix} \varepsilon_t^m \\ \varepsilon_t^y \end{bmatrix}, \quad \alpha_0 = \begin{bmatrix} 1 & -\lambda_{my} \\ -\lambda_{ym} & 1 \end{bmatrix},$$

$$\alpha_j = -\begin{bmatrix} \alpha_{j,mm} & \lambda_{j,my} \\ \lambda_{j,ym} & \lambda_{j,yy} \end{bmatrix}, \quad j = 1, 2, ..., p.$$

Our prime concern here focuses on the long-run dynamic effect of the monetary shock ε_t^m on the real output y_t. We can express this in terms of the long-run multiplier, $\gamma_{ym} = \alpha_{ym}(1)/\alpha_{yy}(1)$, which yields the percentage increase in real output for each percentage point increase in money resulting from a permanent monetary shock. As such, we can also define another long-run multiplier, $\gamma_{my} = \alpha_{my}(1)/\alpha_{mm}(1)$, which is related to the long-run money response to a permanent real output shock. Within this context, the long-run neutrality of money implies the restriction $\gamma_{ym} = 0$.

As King and Watson (1997) cautioned against, the endogeneity of m_t and y_t means that Eq. (3) is econometrically unidentified. This can be clearly seen by writing the reduced form of Eq. (3) as:

$$\boldsymbol{X}_t = \sum_{j=1}^{p} \phi_j \boldsymbol{X}_{t-j} + \boldsymbol{e}_t, \tag{4}$$

where $\phi_j = -\alpha_0^{-1}\alpha_j$ and $\boldsymbol{e}_t = -\alpha_0^{-1}\boldsymbol{\varepsilon}_t$. We determine the matrices α and Σ_ε from the following equations:

$$\alpha_0^{-1}\alpha_j = -\phi_j, \text{ where } j = 1, ..., p, \tag{5}$$

$$\alpha_0^{-1}\Sigma_\varepsilon(\alpha_0^{-1})' = \Sigma_e. \tag{6}$$

Eq. (5) determines α_j as a function of α_0 and ϕ_j. Eq. (6) cannot determine either α_0 or Σ_ε in that Σ_e is a 2×2 symmetric matrix with only three unique elements. Therefore, even under the assumption of the independence of ε_t^m and ε_t^y, we only identify three of the four unknown parameters: λ_{my}, λ_{ym}, var(ε_t^m) and var(ε_t^y). One additional restriction is required in order to identify the model.

To obtain a consistent estimate of γ_{ym}, we adopt King and Watson's (1997) eclectic approach. Rather than focus on a single identifying restriction, we report the results for a wider range of identifying restrictions. To be more precise, we consider three identification schemes that use a pair of restrictions:

1. Σ_ε diagonal, where λ_{my} is known;
2. Σ_ε diagonal, where λ_{ym} is known; and
3. Σ_ε diagonal, where γ_{my} is known.

As emphasized in King and Watson (1997), this testing strategy is unquestionably more informative in terms of the robustness of the inference concerning the link between money and real output to specific assumptions about λ_{ym}, λ_{my} or γ_{my}. we estimate the model using the instrumental variables procedure.[2]

3 Data and Results

3.1 Integration and Cointegration Properties of the Data

Using quarterly frequency of money stock and real output data, we consider two Pacific Basin countries, namely South Korea and Taiwan, in our empirical analysis. The sample size runs from 1970Q1 to 2004Q4 for South Korea and from 1965Q1 to 2004Q4 for Taiwan. We download the data from the *AREMOS* data bank. The traditional augmented Dickey and Fuller (1979, ADF) test results indicate that the null hypothesis of a unit root cannot be rejected for money and real output.[3]

We adopt the Augmented Engle and Granger (1987, AEG) two-step procedure (for which non-cointegration serves as the null hypothesis) and the C_μ statistic from Shin (1994) for which cointegration serves as the null hypothesis. Table 1 presents various cointegration test results. It is most apparent that cointegration between money and real output cannot be rejected by the AEG test at the 5% level. As for the C_μ statistic, the cointegrated relationship between money and real output

[2]Details on the instrumental variables estimation procedure, including the computation of standard errors, are provided in the Appendix to King and Watson (1997). Readers are referred to their paper for details.

[3]We do not show the test results of unit root because of space constraints. However, the results are available from the author upon request.

is rejected at the 5% level. Analogous to the unit root test, we also adopt the AEG^{GLS} cointegration test developed by Perron and Rodriguez (2001). The test is in tune with that of Ng and Perron (2001) but has greater power and size than does the conventional AEG statistic. The AEG^{GLS} test results in Table 2 show that real output is not in fact cointegrated with money stock.

3.2 Permanent Effects

Figure 1 presents the point estimates of γ_{ym} based on the three identification schemes for money stock and real output for South Korea. The dashed lines in each panel delineate the 95% confidence bands for the γ_{ym} estimates. In the first identification condition, presented in the top panel (diagonal Σ_ε and known λ_{my}), the point estimates of γ_{ym} remain stable as the value of λ_{my} increases. For example, the point estimate of γ_{ym} is -4.267 when the value of λ_{my} equals 0 (see Table 2). However, based on the fact that the width of the bands of the confidence intervals do cover zero, they are insignificantly different from zero, indicating that the long-run neutrality of money cannot be rejected.

The results also show that long-run neutrality is not rejected for a reasonable range of values of λ_{ym}. As for the second identification scheme, presented in the middle panel (diagonal Σ_ε and known λ_{ym}), the point estimates of γ_{ym} remain relatively stable for the assumed values of λ_{ym} from -0.5 to 1.5. The point estimate of γ_{ym} is -7.276 when the value of λ_{ym} equals 0. Also, given the wide bands, they are insignificantly different from zero. This is consistent with two traditional monetary models of the business cycle, which implies that $\lambda_{ym} \geq 0$ (that is, output does not decline on impact in response to a monetary expansion), and with Lucas' (1972b) monetary misperceptions theory (due to incomplete information concerning the state of the economy). This suggests that λ_{ym} could be negative (for example, a positive change in the quantity of money could lead to a decrease in output if that change is less than expected). The results are also consistent with the real business cycle model where money is neutral and level changes in money stock have no effect on the real variables and cause a proportionate increase in the price level (Williamson, 2005).

As for the last identification condition (diagonal Σ_ε and known γ_{my}), shown in the bottom panel, the point estimates of γ_{ym} are again relatively unchanged for the assumed γ_{mm} values from -0.15 to 0.12, and they are insignificant. The point estimate of γ_{ym} is 5.533 when the value of γ_{my} equals 0. Clearly, the confidence intervals include many reasonable values of the parameters, and we conclude that they are consistent with long-run money exogeneity. Overall, the three identification conditions provide substantial evidence that the long-run neutrality of money is fully accepted in the case of South Korea.

Figure 2 presents plots similar to those in Figure 1 but for Taiwan. As for the first identification condition, presented in the top panel (diagonal Σ_ε and known λ_{my}), the point estimates of γ_{ym} remains stable as the value of λ_{my} increases. For example, the point estimates of γ_{ym} are 1.8 and 2.643, respectively, when the values of λ_{my} equal 0.3 and 0. However, because the widths of the bands of the confidence intervals do not include zero, they are significantly different from zero. The use of the first identification demonstrates that money has a permanent effect on real output. Our use of the second and third identifications also shows that the widths of the bands of the confidence intervals do not include zero. The point estimates of γ_{ym} are 2.397 and 2.993, respectively, when the values of λ_{ym} γ_{my} equal 0. Thus, similar to the finding with the first identification condition, the long-run neutrality of money is rejected in the case of Taiwan. We conjecture that there are two possible explanations for the fact that long-run monetary neutrality must be rejected in the case of Taiwan. The first is that the definition of monetary aggregate used in this paper is M2 rather than M1. The long-run monetary neutrality hypothesis is also rejected in some cases where different definitions of money stock are employed – for example, Olekalns (1996) for Australia; Coe and Nason (2003) for Australia, Canada, the United kingdom and the United States; Jefferson (1997) for measures of both inside money and outside money; Serletis and Koustas (1998) for the United Kingdom; and Serletis and Koustas (2001) for broadly-defined monetary aggregates. The second explanation is that the sample size used in this paper may not be long enough for long-run monetary neutrality to be satisfied. How much time is required to satisfy long-run monetary neutrality? To date, no exact answer to this question has been given in the literature. It should be noted that the purpose of this paper is not intended to question or overturn the proposition of long-run monetary neutrality in macroeconomics, but just to provide empirical evidence of monetary neutrality in developing economies.

3.3 Short-Run Effect

We also examine the short-run effect of a permanent shock of money on real output by using the impulse response function (IRF), the hallmark of VAR models. Tables 4 and 5 respectively present the estimated IRFs for South Korea and Taiwan using various values.[4] In each case, we impose three identifications, but we only consider one particular set of restrictions. For example, the second column displays the estimated results from the IRF with the restrictions $\lambda_{my} = (0.3, 0.1, -0.5)$; the third column adopts the restrictions $\lambda_{ym} = (0.5, 0, -0.5)$; and the fourth column uses the re-

[4]We do not show the estimated IRF graphs because of space constraints. However, the graphs are available from the author upon request.

strictions $\gamma_{my} = (0.2, 0, -0.1)$.

In the case of South Korea (Table 3), the results from the IRF can be summarized as follows. With the first identification, the permanent shock of money has a positive effect on real output when the assumed value of λ_{my} is -0.5, but it has a negative effect when the assumed values of λ_{my} are 0.3 and 0.1. The short-run effect of a monetary shock on real output is negative when the assumed values of λ_{ym} are 0.5, 0, and -0.5. The IRFs of a monetary shock on real output are positive when γ_{my} is assumed to be 0 and -0.1 but are negative when the assumed value of γ_{my} is 0.2. The estimated IRF results of the three identifications show that monetary shocks have a permanent effect on real output in the short-run in the case of South Korea.

In the case of Taiwan (Table 4), with the first identification, a permanent shock of money has a positive effect on real output when the assumed values of λ_{my} are -0.5 and 0.1, but it has a negative effect when the assumed value of λ_{my} is 0.3. With the second and third identifications, i.e., λ_{ym} and γ_{my}, the IRFs of a monetary shock on real output are all positive when $\lambda_{ym} = 0.5$, 0, -0.5 and $\gamma_{my} = 0.2$, 0, -0.1. The short-run IRFs confirm that monetary shocks have a real effect on real output. Again, empirical evidence shows that the short-run neutrality of money in the case of Taiwan is rejected.

4 Concluding Remarks

We examine the long-run and short-run neutrality of money for South Korea and Taiwan. We test the long-run as well as the short-run real output response to a permanent monetary shock using King and Watson's (1997) eclectic approach. The empirical evidence shows that the long-run neutrality of money is fully supported in the case of South Korea, but not in the case of Taiwan. This is consistent with two traditional monetary models of the business cycle, that is, the monetary intertemporal model and Lucas' (1972b) monetary misperceptions theory. The results are also consistent with the real business cycle model where money is neutral, i.e., level changes in money stock have no effect on the real variables and cause a proportionate increase in price level. The estimated results from the IRFs indicate that the hypothesis of the short-run neutrality of money must be rejected for South Korea and Taiwan.

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Table 1: Cointegration Test

Country	AEG^{GLS} statistic		C_{μ} statistic	
	M2	*Real Output*	*M2*	*Real Output*
South Korea	-1.858	-1.770	0.309**	0.305**
Taiwan	-1.578	-1.558	0.463*	0.248**

* denotes significance at the 5% level.

** denotes significance the 10% level.

Table 2: Estimated Results of γ_{ym}

Country	Point estimate of γ_{ym}		
	$\lambda_{my} = 0$	$\lambda_{ym} = 0$	$\gamma_{my} = 0$
South Korea	-4.267	-7.276	5.533
Taiwan	2.643	2.397	2.993
	Confidence interval estimates of γ_{ym}		
	$\lambda_{my} = 0$	$\lambda_{ym} = 0$	$\gamma_{my} = 0$
South Korea	[-19.82, 11.29]	[-34.71, 20.16]	[-18, 29.07]
Taiwan	[1.301, 3.985]	[0.885, 3.910]	[1.645, 4.340]

Table 3: Estimates of Short-Run Impulse Response — South Korea

Monetary shock to real output									
	λ_{my}			λ_{ym}			γ_{my}		
period	0.3	0.1	−0.5	0.5	0	−0.5	0.2	0	−0.1
1	−4.7	1.0	−3.8	−1.5	1.0	3.0	5.2	−5.0	−7.0
5	−9.5	4.2	−1.2	1.5	4.0	6.0	7.7	−3.0	−6.5
11	−7.8	3.9	−0.3	1.5	3.5	5.0	6.0	−2.0	−5.0
20	−8.0	4.0	−0.5	1.8	3.8	5.3	6.5	−2.0	−5.0
Real output shock to money									
	λ_{my}			λ_{ym}			γ_{my}		
period	0.3	0.1	−0.5	0.5	0	−0.5	0.2	0	−0.1
1	−25.0	0	7.0	3.2	0	−3.0	−7.0	8.0	10.0
5	−39.0	−3.0	8.0	1.2	−4.0	−8.0	−13.0	10.0	15.0
11	−55.0	−8.0	9.0	0	−7.5	−14.0	−20.0	13.0	25.0
20	−83.0	−10.0	12.0	−0.5	−12.0	−21.5	−30.0	20.0	32.0

Table 4: Estimates of Short-Run Impulse Response — Taiwan

Monetary shock to real output									
	λ_{my}			λ_{ym}			γ_{my}		
period	0.3	0.1	−0.5	0.5	0	−0.5	0.2	0	−0.1
1	−4.0	0.8	−2.0	0	0.8	1.5	−0.6	−1.5	−1.6
5	−8.0	0.9	−5.0	−0.5	1.7	3.0	−1.8	−4.0	−4.4
11	−10.0	2.0	−4.0	0.8	2.8	4.2	−0.9	−3.2	−3.5
20	−16.0	4.5	−4.0	2.1	4.8	7.0	0.3	−2.7	−3.3
Real output shock to money									
	λ_{my}			λ_{ym}			γ_{my}		
period	0.3	0.1	−0.5	0.5	0	−0.5	0.2	0	−0.1
1	−10.0	0	5.0	1.8	0	−1.2	2.2	3.8	3.8
5	−17.0	5.0	10.0	6.0	3.5	0.8	7.3	9.0	9.0
11	−25.0	16.0	8.0	2.8	6.8	10.2	13.0	15.9	16.0
20	−35.0	14.0	25.0	18.0	12.8	6.8	21.2	24.8	25.2

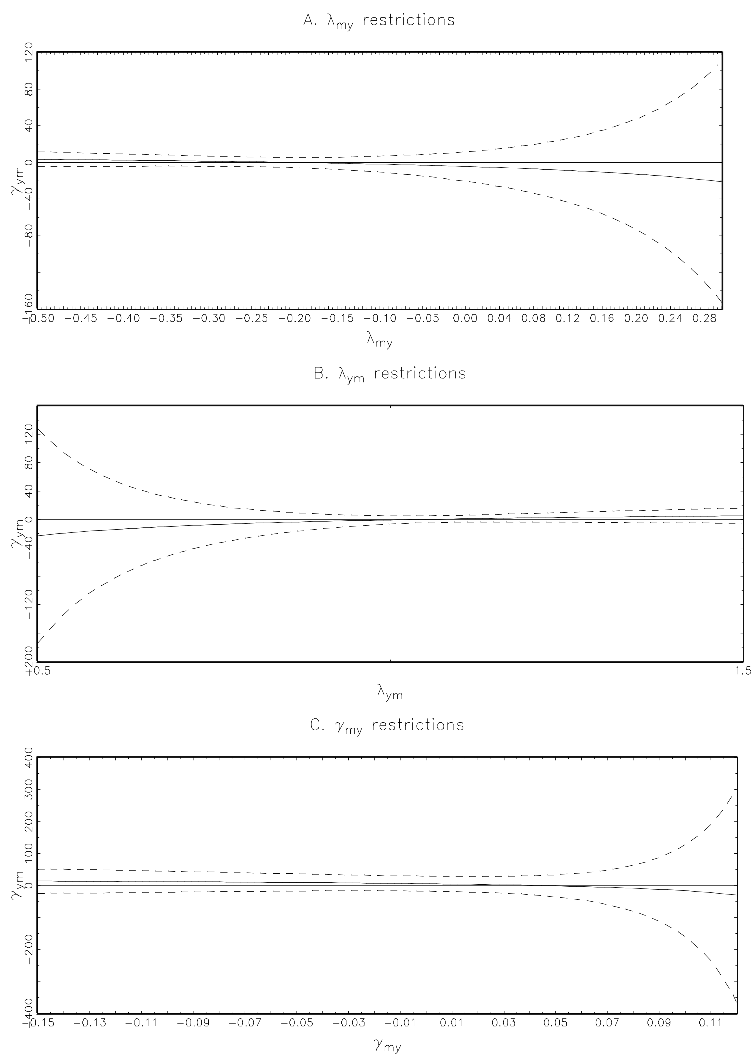

Figure 1: γ_{ym} estimates for different λ_{my}, λ_{ym} and γ_{my} — South Korea.

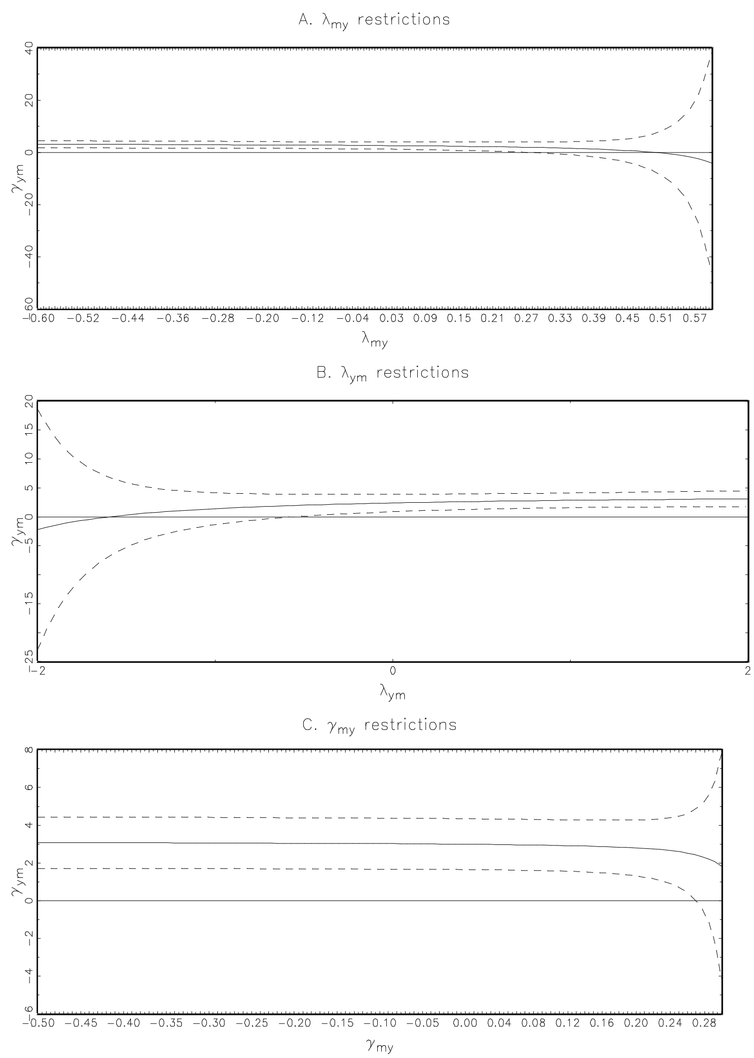

Figure 2: γ_{ym} estimates for different λ_{my}, λ_{ym} and γ_{my} — Taiwan.