CROWDING-IN OR CROWDING-OUT? ANALYZING GOVERNMENT INVESTMENT IN TAIWAN*

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Economic theory suggests that the way government finances its expenditure determines the effectiveness of fiscal expansionary policy. This paper considers a simple investment model embedded in a Markov regime-switching framework, where parameters are subject to shift between two regimes: crowding-in and crowding-out (of private investment by government investment). Using Taiwanese data, the study finds dominant crowding-in effects before 1980, and dominant crowding-out effects after 1980. The dating correctly separates two exchange rate regimes (a fixed, then a flexible rate regime). One conclusion is that fiscal policy is ineffective in a flexible-rate regime.

INTRODUCTION

The threat of fiscal crowding-out usually renders unconvincing the plea for expansionary fiscal policy. The conventional negative relationship between private investment and government investment, which is called crowding-out, can be used to evaluate the fiscal effectiveness of expansionary public investment. Empirical evidence for the crowding-out is quite mixed. On the one hand, first Ahmed (1986) estimated the effects of the United Kingdom (UK) government consumption in an inter-temporal substitution model and found that government expenditure tends to crowd out private consumption. Recently, Aiyagari *et al.* (1992) and Baxter and King (1993) explored the effect of government spending shocks on various economic aggregates in a one-sector neoclassical growth model with constant returns to scale and variable labour supply. They found that increases in government spending significantly lead to a decline in private consumption. Amano and Wirjanto (1997) tested this hypothesis by estimating the intra-temporal substitution elasticity to be about 0.9, assuming that a representative consumer maximizes his or her lifetime utility by consuming two goods.

On the other hand, some empirical studies have found different results. In terms of a neoclassical model with increasing returns to scale and monopolistic competition, Devereux, Head and Lapham (1996) examined the impact of government spending shocks and found that an increase in government consumption generates an endogenous rise in aggregate productivity. The increase in productivity raises the real wage sufficiently that there is a substitution away from leisure and into consumption. Thus, an increase in government expenditures leads to an increase in private consumption. Karras (1994) examined the change of private consumption in response to increases in government spending across a number of countries and found that public and private consumption are better described as complementary rather than as substitutes.

There are two reasons motivating this study. First, over a long sample period, it is implausible to argue that the extent of crowding-out or crowding-in is invariable. Economic theories suggest one possibility that may change this relationship, for example, is the exchange rate regimes in the Mundell-Fleming model. Thus, it is reasonable to assume that an economy switches between crowding-out and crowding-in regimes over time.

Second, total government expenditure is too aggregate to permit any inference of plausible explanation for fiscal policy effectiveness; for instance, government spending can be broken into several budget categories. It is unreasonable to argue that education and other welfare expenditure will be less effective in stimulating the economy on the basis that their budgetary purposes are irrelevant to economic stimuli. For this reason, public investment is used here as an alternative instrument to investigate the effect of crowding-out.

Taiwan is a good case for empirically examining this problem. First, its private sector was heavily dependent on government investment in its initial stage of economic development; that is, government investment might crowd in the private investment in the beginning, but it may be shown to crowd out private investment in the later stage of economic development. Second, during recent decades, Taiwan has experienced a structural financial reform and the private sector has grown up; that is, the effect of government investment on private investment now may not be as beneficial as previously. Therefore, at any point of time, both crowding-in and crowding-out effects might co-exist. What is important is to determine which effect dominates over time.

Moreover, the economic development of Taiwan has been based on the active role of government investment from the beginning; however, due to a sequence of institutional changes, the role of government may not be as effective now. This paper attempts to investigate the changing relationship between government investment and private investment in Taiwan over time.

Due to data availability, quarterly data from the first quarter of 1961 to the last quarter of 1999 are derived from the *Taiwan Statistical Data Book*, published annually by the Council for Economic Planning and Development of the Executive Yuan of Taiwan, Republic of China. Table 1 (p. 76) offers the summary statistics. This paper is organized as follows. After giving a rundown on financial reforms and public investment in Taiwan, the study develops the empirical framework and then empirical results are presented. The next part proposes a Markov-switching modelling framework and its policy implications of estimation results are discussed. Finally, there is a concluding section.

TABLE 1
Summary statistics
(NT$ million)

	Private Investment	Government Investment
Mean	61676.500	26651.260
Median	51098.500	16439.500
Maximum	256663.000	148707.000
Minimum	3377.000	503.000
Std. Dev.	57048.050	32969.690
Skewness	1.151	1.727
Kurtosis	3.896	5.253
Jarque-Bera (*p*-value)	29.48 (0.000)	82.160 (0.000)

Note: The Jarque-Bera statistic tests the null of normality.

TAIWAN'S FINANCIAL REFORM AND PUBLIC INVESTMENT

Up to the mid-1980s, more than 50% of funds of households and non-profit institutions was channelled to regulated financial market intermediaries, notably banks. The saving ratio was high, while the interest rates were kept at a low level with a view to providing low-cost funding for policy-designated key industries. The financial market was quite shallow and under-developed. During the period from 1980 to 1985, enterprises listed on the stock market accounted for only 16% of their funds to the stock market during that period. Financial sector regulations tended to segment financial markets and reduced the allocation of funds out of the regulated financial sectors. After 1989, barriers to entry and interest rate limits were deregulated, and restrictions on branching were also gradually eased.

Taiwan has gradually liberalized its interest rates since 1980. More precisely, the central bank deregulated the discount rates on certificates of deposit (CDs) and bank debentures and allowed the Banker's Association to set the range of the maximum and minimum lending rates. In September 1984, the central bank informed individual banks that they could set their own prime rate based on their cost of funds; Figure 1 (p. 87) indicates that the prime lending rate began to decline and tended to move toward a specific range over time. In January 1986, the deposit interest rate ceilings were simplified, with a reduction from twelve rates to four, and the spread of maxi–mini lending rates was enlarged. Figures 2 and 3 (p. 87) illustrate this, showing that the one-month and one-year deposit rates have exhibited larger fluctuations since then. Moreover, Figure 4 (p. 87) also indicates that the over-night interbank rate first declined and then fluctuated around a lower mean. These plots provide evidence that the market force has been affecting the interest rate. Finally, in July 1989, the revised banking law lifted all restrictions on interest rates, bringing the long history of interest rate control in Taiwan to an end. The reason for the continuous loosening of interest rate control after 1985 was the pressure from the fast-growing money supply and the high saving ratio that flooded banks with excess liquidity. The banking rate was forced to adjust, in line with the money market interest rates. However, it must be emphasized that, while price liberalization has made considerable progress, it would still be premature to say that the market in Taiwan principally determines interest rates. The government continues to have a dominant influence over financial sectors. Entry barriers still exist in the oligopolistic structure of the banking industry, although the privatization of state-owned banks and market entry have been permitted.

THE FRAMEWORK OF EMPIRICAL ANALYSIS

The Theoretical Model

Aggregate investment is an important component of the standard national account. Items of government investment are different from those of private investment; for instance, they may be related to infrastructure spending and state-owned enterprises' investments, so they may exert different effects on the private investment. Hence, effective investment function is defined here as the aggregate investment level that effectively utilizes production capacity and attains necessary economic growth. The effective investment I_t^* is specified as

$$I_t^* = I_t + \beta \cdot GI_t \tag{1}$$

where I_t is the real private investment, GI_t is the real government investment at period t, and β is a time-varying parameter measuring the effect of government investment on effective investment over time. A representative production sector chooses an effective investment level to maximize expected lifetime investment utility function U:

$$\text{Max } E_0\left[\sum_{t=0}^{\infty} \rho^t U\left(I_t^*\right)\right] \tag{2}$$

$$\text{s.t } A_{t+1} = \left(A_t + Y_t - C_t - I_t^* - (1-\alpha)\cdot GI_t\right)\cdot(1+r) \tag{3}$$

where E_t is the expectations operator based on information of period t. ρ is a discount factor. Equation (3) is the budget constraint, where A_t is the real financial assets net real government debt at the beginning of period t and r is a time invariant real rate of interest. Finally, it is assumed that U is increasing and concave in its arguments and that $\partial U(0)/\partial I^* \to \infty$. As in Barro (1989) and Christiano and Eichenbaum (1992), a function of GI can be added to the utility function so that the government investment's marginal utility becomes positive. Hence, equation (2) can be written below as:

$$U_0 = E_0\left[\sum_{t=0}^{\infty} \beta^t \left\{U\left(I_t^*\right) + \left(GI_t\right)\right\}\right] \tag{4}$$

with $\partial\phi/\partial G > 0$; and under the usual assumption that the agent (production sector) has no control over government investment spending (GI), the optimization problem can be solved, ignoring the government investment's contribution to the objective utility function through ϕ. Accordingly, the agent's problem is now to maximize equation (4) subject to equation (2). It can be easily shown that the optimal sequence of I_t^* must satisfy the Euler equation below:

$$\frac{U'\left(I_t^*\right)}{\rho \cdot E_t U'\left(I_{t+1}^*\right)} = 1 + r \tag{5}$$

Equation (5) simply says that, along the optimal path, the inter-temporal rate of substitution must be equal to the inter-temporal rate of transformation. In the next section, the restriction imposed by equation (5) will be utilized in order to estimate the parameter β. To investigate the empirical implications of the model, this study employs the famous Hall's (1978) result that the marginal utility of consumption follows a random walk, so that equation (5) can be written as:

$E_t I^*_{t+1} = [\rho(1+r)]^\sigma \cdot I^*_t$, where $\sigma = \left(-U'(I^*)/\{I^* \cdot U''(I^*)\}\right)$ is the inter-temporal elasticity of substitution, which also represents a coefficient of risk aversion and is assumed to be constant over time (see Hall 1978, for details). Therefore, the econometric relationship is approximated below:

$$I^*_{t+1} = \gamma \cdot I^*_t + v_{t+1} \tag{6}$$

where $E_t[v_t] = 0$. Accordingly, equations (1) and (6) imply a model below:

$$I_t + \beta \cdot GI_t = \gamma \cdot (I_{t-1} + \beta \cdot GI_{t-1}) + v_t \tag{7}$$

To generalize, the empirical form of equation (7) can be specified with intercept below:

$$(I_t - \gamma \cdot I_{t-1}) = \alpha + \beta \cdot (GI_t - \gamma \cdot GI_{t-1}) + v_t \tag{8}$$

If $\gamma = 1$, it implies the effective investment I^*_t has a unit-root; that is, both I_t and GI_t are integrated of order 1 and are cointegrated in the sense of Engle and Granger (1987) with cointegrating vector (α, β). Equation (8) hence implies an underlying error correction mechanism, which can be consistently estimated by the procedures suggested by Wickens and Breusch (1988), Phillips and Hansen (1990), Phillips and Loretan (1991) and Park (1992). The following sections continue the empirical analysis.

Analysis of Time Series Properties

First, Table 2 (opposite) reports the unit root tests for both variables as well as their first-order difference. Before they are formally tested, Figure 5 (p. 88) plots them and visually confirms the presence of unit root. To formally test for the presence of unit roots, Augmented Dickey-Fuller (ADF) (Said & Dickey 1984) and KPSS (Kwiatkowski *et al.* 1992) tests are employed. The KPSS tests the null [hypothesis] that the series is stationary in trend; and, at the 5% significance level, both variables reject the null. The ADF statistic tests the null that the series is non-stationary. At the 5% significance level, critical value is –3.46. Hence, for both variables, the null cannot be rejected. These tests unanimously confirm the presence of unit roots. Banerjee *et al.* (1986), Phillips (1987, 1991) and Phillips and Ouliaris (1990) have shown that conventional tests in multivariate regressions with integrated processes cannot be applied asymptotically. In this case, classical asymmetric theory breaks down and the presence of nuisance parameter dependencies in the limiting distribution theory raises similar issues to panel data with $I(1)$ processes. To expose this problem, assuming that the generating mechanism for Y_t is a cointegrating system

$$Y_t = \alpha + \beta X_t + u_{1t} \quad t = 1,2,\ldots T \tag{9}$$

$$\Delta X_T = u_{2t} \tag{10}$$

Phillips and Durlauf (1986) showed that, under appropriate centering and scaling, Ordinary Least Squares (OLS) estimation of the cointegrating vector in equation (9) is asymptotically non-normal. The weak convergence of appropriately scaled sample moments to random matrices rather than constant matrices results in this non-normality. Moreover, OLS leads to estimators

TABLE 2
Tests for unit root

Variables	KPSS		ADF
	Level (0.461)	Trend (0.146)	(-3.46)
I	0.978	0.223	-2.46
GI	0.978	0.226	-1.58
Prime	0.822	0.393	-2.27

Notes: Numbers in the parentheses are critical values at the 5% significance level. Models are estimated by specifying 12 lags, along with trend and intercept.

that are asymptotically biased and whose distributions involve unit root asymptotics and non-trivial nuisance parameters (see Phillips & Loretan 1991: 426). Phillips (1987, 1991) and Phillips and Ouliaris (1990) have proved that standard tests statistics, such as the Wald test, no longer generate asymptotically distributed χ^2 criteria. Because the limiting distribution of the regression coefficients is non-normal, the metric underlying the Wald test is no longer valid. In other words, statistical estimation and inference in these models require a methodology that accounts for the non-stationarity of the underlying time series. First, let us consider the univariate error correction model (ECM) below,

$$\Delta I_t = \alpha + \beta \cdot \Delta GI_t + \delta \cdot \left(I_{t-1} - GI_{t-1}\right) + \varepsilon_t \qquad (11)$$

The size of δ is a measure of the sustainability of investment gap, which has a natural interpretation: it measures the speed of convergence of the system toward equilibrium. Moreover, following the test outlines by Kremers, Ericsson and Dolado (1992), a hypothesis test of δ being non-zero may be interpreted as equivalent to a test of cointegration. Converted into a 'half-life' measure, this offers an indication of the sustainability of current-account disequilibria for our panel in the sample period under study. To make it simple, parameters estimates will not be examined now; they will be examined later, using vector error correction model (VECM). Here, an estimate is made to check equation (11) for stability and the recursive residuals for model stability are plotted in Figure 6 (p. 88). Evaluated by ± 2 standard errors, the analysis indicates that the system is relatively unstable.

Johansen's (1988, 1990) VECM is a multivariate model, which considers a general vector autoregression (VAR) model with Gaussian errors expressed in the error correction form

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \beta X_{t-k} + \Phi D_t + \mu + \varepsilon_t \qquad (12)$$

where $t = 1, 2, \ldots, T, X_t$ is p-dimensional vector, D_t are seasonal dummies orthogonal to the constant terms and $\varepsilon_t \sim N(0, \Lambda)$. The null hypothesis of non-stationarity is determined by whether the rank of β is full. If Rank $(\beta) = p$, then it is full rank and the model is cointegrated. Johansen's likelihood ratio test statistic is

$$-T \sum_{i=r+1}^{p} \ln\left(1 - \lambda_i\right) \qquad (13)$$

which can be obtained by first running two regressions below:

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_{0i} \Delta X_{t-i} + R_{0t} \qquad (14)$$

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_{ki} \Delta X_{t-i} + R_{kt} \qquad (15)$$

The second step is to solve the eigenvalues from the equation $\left| \lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k} \right| = 0$, where

$$S_{ij} = \frac{\sum_{t=1}^{T} R_{it} R'_{jt}}{T} \text{ and } i, j = 0, 1, \ldots, K.$$

The empirical analysis here is divided into two parts: the first part fits the vector $[I, GI]$ of equation (8) into VECM; the second part fits the vector $[I, r]$ into VECM where r denotes prime lending (loan) rate and GI is introduced as an exogenous variable. The second part aims at examining the interest rate effect of exogenous government investment on private investment. To compare the empirical results, regressions are conducted over three sampling periods: 1961Q1–1980Q1; 1980Q2–1999Q4; and full sample.

Table 3 (opposite) presents the results estimated using vector $[I, GI]$, which has two conclusions: first, they indicate a crowding-in effect of government investment on private investment, due to the positive and significant estimate β over the three sampling periods; second, the tests for cointegration reject the no-cointegration null, in favour of cointegration. Figure 7a (p. 88) plots the residuals for two series. It shows that private investment has larger disturbance than government investment, especially in the post-1985 period.

Table 4 (p. 82) presents the results using vector $[I, r]$. The prime lending rate also has unit roots (see Table 2, p. 79). In addition to the estimation results of three sub-sample periods, the vector is divided into two specifications: the first specification estimates the vector $[I, r]$ incorporating exogenous government investment; the second specification estimates the vector $[I, r]$ excluding the exogenous government investment. The authors note two important findings: first, most parameter estimates are insignificant when the government investment is excluded; second, the parameter estimates of three sub-samples exhibit different results for two specifications; while they show consistent results when the full sample is used. Figure 7b (p. 89) plots the residuals for this specification. Private investment still fluctuates out of bounds for most of the time during the post-1975 period, implying a greater volatility as shown in the previous study (Figure 7a, p. 88); however, it may be not informative enough to say so.

In general, the study shows mixed results concerning the impact of exogenous government investment on private investment via the interest rate channel. In a nutshell, it implies that the government investment may have regime-switching effects on private investment: crowding-in and crowding-out. Accordingly, the above-mentioned results motivate a subsequent analysis, using Markov-switching modelling. Under Markov regime-switching framework, the parameters are subject to change by states of unobserved Markov chain. Hence, in the Markov regime-switching model, at any point of time, there is probability for the existence of both states: crowding-in and crowding-out; and the timing of each state can be identified afterwards, which enables the study to examine which state dominates.

Conventional econometric inferences are heavily dependent upon the untested underlying assumptions of time-invariance of time-series process with no changes in structure (financial reform). Even so, regime changes and structural breaks are both economically and empirically relevant and can severely affect the properties of inferential procedures. In light of this,

TABLE 3
Results from VECM: private investment and government investment

Variables	1961:1–1980:1	1980:2–1999:4	Full Sample
Intercept	3.11 (0.46)	5.56 (0.58)	2.36 (0.48)
Government investment	0.75 (0.05)	0.54 (0.06)	0.83 (0.04)
AIC	-0.84	-2.83	-1.29
SBC	-0.36	-2.38	-0.98
Cointegration test			
*H*o: Number of cointegrating equations	LR Statistic	LR Statistic	LR Statistic
None	32.04** (19.96)	24.29* (19.96)	49.01**(19.96)
At most one	6.84 (9.27)	5.04 (9.27)	10.41*(9.27)

** and * indicate significance at the 1% and 5% significance levels, respectively.

Notes: Numbers in the parentheses of parameter estimates are standard errors. Numbers in the parentheses of LR (Likelihood Ratio) statistics are critical values at the 5% significance level under the null.

a regime-switching approach is employed to examine this issue empirically. This distinguishes the study from others in one important aspect: it explicitly takes into account the regime-switching properties of the parameter where the relationship between government investment and private investment is subject to transition probabilities. Unlike conventional structural change models, the regime-switching model does not exclude any effect from either exchange rate regime. Instead, it examines the relative persistence of each effect in the whole sample period. The next sub-section explains.

Markov-switching ECM model

Hamilton (1989) pioneers the work on the Markov-switching model. To account for non-stationarity, equation (11) is embedded in a simple Markov-switching ECM (MSECM) below,

$$\begin{aligned}\Delta I_t =& \left[\alpha_1 \cdot (1-S_t) + \alpha_2 \cdot S_t\right] \\ &+\left[\beta_1 \cdot (1-S_t) + \beta_2 \cdot S_t\right] \cdot \Delta GI_t \\ &+\left[\delta_1 \cdot (1-S_t) + \delta_2 \cdot S_t\right](I_{t-1} - GI_{t-1}) \\ &+\left[\sigma_1 \cdot (1-S_t) + \sigma_2 \cdot S_t\right] \cdot \varepsilon_t\end{aligned} \quad (16)$$

where $\varepsilon_t \sim i.i.d.N(0,\sigma^2)$ and with unobserved state S_t, which is presumed to follow a two-state Markov chain with transition probabilities p_{ij}. Negative δs represent cointegration. Signs of two slope parameters (β_1, β_2) represent two cointegrating regimes: crowding-in and crowding-out. The evolution of the unobservable state variable is assumed to follow a two-state Markov chain satisfying: $p_{11} + p_{12} = p_{22} + p_{21} = 1$, where $p_{11} = \Pr(S_t = 1 | S_{t-1} = 1)$, $p_{12} = \Pr(S_t = 2 | S_{t-1} = 1)$, $p_{21} = \Pr(S_t = 1 | S_{t-1} = 2)$, $p_{22} = \Pr(S_t = 2 | S_{t-1} = 2)$. The observation can also be thought of as drawing from a mixture of two normal distributions. The state in each period determines which of the two normal densities is used to generate the model. Their correlation is

TABLE 4
Results from VECM: private investment and prime loan rate

Variables	1961:1–1980:1	1980:2–1999:4	Full sample
	Including exogenous government investment		
Intercept	8.47 (8.53)	6.24 (0.25)	5.21 (0.83)
Prime rate	4.44 (3.13)	-0.196 (0.05)	-0.51 (0.22)
AIC	-3.91	-3.42	-3.41
SBC	-3.13	-2.88	-3.05
Cointegration test			
*H*o: Number of cointegrating equations	LR statistic	LR statistic	LR statistic
None	17.54 (19.96)	57.49** (19.96)	30.82**(19.96)
At most one	5.41 (9.27)	7.44 (9.27)	6.91 (9.27)
	Excluding exogenous government investment		
Intercept	104.85 (158.8)	31.8 (203.5)	21.9 (3.81)
Prime rate	-35.51 (58.36)	19.8 (93.16)	-4.28 (1.52)
AIC	-3.82	-2.68	-3.28
SBC	-3.36	-2.61	-2.97
Cointegration test			
*H*o: Number of cointegrating equations	LR statistic	LR statistic	LR statistic
None	23.74* (19.96)	12.25 (19.96)	22.3* (19.96)
At most one	7.65 (9.27)	4.71 (9.27)	6.56 (9.27)

** and * indicate significance at the 1% and 5% significance levels respectively.

Notes: Numbers in the parentheses of parameter estimates are standard errors. Numbers in the parentheses of LR (Likelihood Ratio) statistics are critical values at the 5% significance level under the null.

assumed to switch between two regimes according to transition probabilities. For example, when the current correlation relationship is in state 1, there is p_{11} chance for the next correlation to stay in the same regime; when the current correlation relationship is in state 2, there is p_{22} chance for the next correlation relationship to stay in the same regime. The estimation procedure begins with the unconditional probability of the state of the first observation suggested by Hamilton (1994); and Kim's (1994) recursive algorithm is also used for weighting data. The specification test concerned here is

$$H_0: \alpha_1 = \alpha_2, \beta_1 = \beta_2, \delta_1 = \delta_2 \text{ and } \sigma_1 = \sigma_2 \qquad (17)$$

White (1987) presents a general score-based test for mis-specification in maximum likelihood models that leads to several immediately useful tests in the switching regression context considered here. Hamilton (1996) discusses these tests in the context of the Markov mixture of normal distributions model and presents evidence that White's tests tend to over-reject the null in small samples. Accordingly, all the test statistics presented below are interpreted using the (1%) significance level as Hamilton suggested. Given a likelihood function $L(y_t \mid x_t, \theta)$ mentioned previously, $h_t(\theta)$ is simply the gradient of $L(y_t \mid x_t, \theta)$ with respect to θ. White (1987) constructs the general test by listing those l elements of $m \times m$ matrix $h_t(\theta) \times h_t(\theta)'$ that these authors wish to test in the $l \times 1$ vector $c_t(\theta)$. White then lets $\hat{\theta}$ denote the maximum-likelihood estimate of θ and lets $\hat{A}$ be the 2×2 sub-block of the inverse of the partitioned matrix below. Let A denote the matrix below

$$\frac{\begin{bmatrix} \sum_{t=1}^{T} h_t(\theta) \cdot h_t(\theta)' & \sum_{t=1}^{T} h_t(\theta) \cdot c_t(\theta)' \\ \sum_{t=1}^{T} c_t(\theta) \cdot h_t(\theta)' & \sum_{t=1}^{T} c_t(\theta) \cdot c_t(\theta)' \end{bmatrix}}{T} \tag{18}$$

where T is the sample size. In this case, White shows that if the model is well specified, the matrix product

$$\frac{\left[\sum_{t=1}^{T} c_t(\hat{\theta})\right]' \cdot A \cdot \left[\sum_{t=1}^{T} c_t(\hat{\theta})\right]}{T} \tag{19}$$

will have a $\chi^2(l)$ asymptotic distribution. The tests of parameters of two regimes follow the conventional Wald test statistic formula. The results presented below test for three kinds of mis-specification: omitted serial correlation; omitted heteroscedasticity; and Markov state-dependence. Specifically, first-order serial correlation in the derivative of the likelihood function with respect to both intercepts terms would indicate the presence of an AR(1) error process in both regimes. The intuition here is that serial correlation in these gradients implies that there is a tendency to find 'runs' where the constant should be higher or lower, which in this context implies persistence in the residuals, or serial correlation. Similar correlations in the derivatives, with respect to the standard deviation (σ) of both regimes, amounts to tests for first-order regime-specific autoregressive conditional heteroscedasticity (ARCH) effects, since persistence here implies that the volatility in each regime seems to vary over time in a way captured by a first-order autoregression; so it would indicate the presence of ARCH(1) effects in their respective regimes. The presence of such first-order serial correlation in the derivative with respect to transition probability would be evidence of state-dependence in the classification probabilities and implies that a Markov-switching regression would be more appropriate. Interested readers are referred to Norden and Vigfusson (1996).

Estimation Results of MSECM and Discussions

Because of the scaling problem of the intercept, each variable is scaled by multiplying 0.001 to reduce the complexity in computing the intercept. Table 5 (p. 85) reports the maximum likelihood estimates of equation (16). Two regimes are identified, which are interpreted as crowding-out and crowding-in, respectively. The null hypothesis that $H_0: \alpha_1 = \alpha_2, \beta_1 = \beta_2, \delta_1 = \delta_2$ and $\sigma_1 = \sigma_2$ tests whether there are two regimes. The Wald statistic is 127.1, which is distributed as Chi-square distribution with eight degrees of freedom. It rejects the null, indicating the two regimes are significantly identified. Moreover, both autoregressive (AR) and ARCH tests of the lower panel confirm that the model fits the data pretty well, according to the 1% significance level.

The crowding-out regime indicates that $1 increase in government spending would crowd out $0.422 private investment. The second regime indicates that $1 increase in government spending would crowd in \$0.721 private investment. The transition probability indicates that, if the state of current period is in the first regime, there is 91% chance that it would stay in the same state in the next period, so that this regime will persist on average for 1/(1-0.91)=11 quarters. And if the state of current period is in the second regime, there is 86% chance that it would stay in the same state in the next period, which will typically persist for 1/(1-0.86)=7 quarters. The transition probabilities do not inform us which regime dominates over time.

In addition, although transition probabilities show that both regimes are likely to have a persistent effect over time, the inferred transition probabilities plot gives more information, as illustrated in Figure 8 (p. 89). This clearly indicates that Taiwan's government investment has had a stronger crowding-out effect on private investment since 1980; the government investment had a stronger crowding-in effect on private investment prior to 1980.

There are two possible explanations (see Shirley (1983) and Gustav (1992) for relevant discussions) for the timing of crowding-out regime. First, the multiplier process means that an increase in government spending, or any other exogenous increase in spending, produces a greater ultimate effect on the nominal level of income through price increases, real income increases, or both, depending on the state of the economy relative to full employment. Although the authors have no empirical evidence to show whether the 1990s is a high-employment era, it is known that Taiwan exhibits a strong consumption propensity, and had a high saving rate in the 1990s (see Table 6, p. 86).

The second explanation is related to the famous Mundell-Fleming theory. It argues that, given imperfect capital mobility, the ability of expansionary fiscal policy to affect aggregate demand is related to the flexibility of the exchange rate. Under flexible exchange rate regimes, expansionary fiscal policy can stimulate aggregate demand, but it also raises the interest rate which is supposed to reduce domestic private investment. The intuition behind these theoretical results is straightforward. A stronger version of this perspective was first documented by Mundell and Fleming in the early 1960s; and McKibbin and Sachs (1991) developed a dynamic version of the Mundell-Fleming approach, concluding that, given a floating exchange rate, commercial policy would be less effective in improving the United States (US) trade deficit.

During the past decades, the foreign exchange market of Taiwan has experienced a structural transition from a fixed to a flexible regime. Between 1969 and 1978, the New Taiwan (NT) dollar exchange rate was pegged to the US dollar. The exchange rate was influenced by policy-makers trying to preserve the competitiveness of the export sector. Before being forced to change to the regular bid and offer trading mechanism, fluctuation in the exchange rate was limited to 2.25% from the average central rate of the previous trading day, which induced a persistent expectation of NT dollar depreciation. On 10 October 1978, Taiwan announced that it would terminate the fixed exchange rate, and the foreign exchange market began to operate formally on 1 February 1979. Since then, following the establishment of the foreign exchange market, the exchange rate (that is, the central rate) has been determined by the market forces of supply and demand, based on the transactions of the previous day. The foreign exchange control has been largely loosened since 1987 and the mid-rate trading system was abandoned in the foreign exchange market in 1989. In April 1989, the central rate was also abandoned and a relatively flexible foreign exchange system has been in operation ever since. Figure 9 (p. 89) plots the exchange rate movement over this period. At the same time, the restrictions pertaining to the capital account have been gradually relaxed. Until 1987, capital control curbed

TABLE 5
Estimation results of MSECM

	α	β	δ	σ
Regime 1	1.786 (0.352)	-0.422 (0.137)	-0.172 (0.047)	1.349 (0.25)
Regime 2	1.116 (0.053)	0.721 (0.051)	-0.491 (0.089)	0.287 (0.039)

P_{11} = 0.91
P_{22} = 0.86
AIC = -13.76

Specification tests

Null hypotheses	Wald statistics
$H_0: \alpha_1 = \alpha_2, \beta_1 = \beta_2, \delta_1 = \delta_2$ and $\sigma_1 = \sigma_2$	127.1 $\sim \chi^2(8)$
H_0: There is AR(1) effect in Regime 1	2.21
H_0: There is AR(1) effect in Regime 2	2.61
H_0: There is ARCH(1) effect in Regime 1	3.89
H_0: There is ARCH(1) effect in Regime 2	3.62
H_0: There is higher-order Markov effects in Regime 1	4.14
H_0: There is higher-order Markov effects in Regime 2	3.12

Notes: Numbers in the parentheses are standard errors. Critical values of $\chi^2(1)$ with the significance level of 10%, 5% and 1% are, respectively, 2.71, 3.84 and 6.63. As suggested by Hamilton (1992), the 1% significance level is used to evaluate AR(1) and ARCH(1).

capital outflows but did not effectively control capital inflows. Subsequently, the policy turned around, with liberalized capital outflows and restricted inflows, but the new policy was equally ineffective. The policy effect on capital account is plotted in Figure 10 (p. 90); which, together with Table 5, explains the hypothesis of fiscal policy ineffectiveness during a flexible exchange rate regime.

There are other possible explanations for the occurrence then of the crowding-out regime. First, government investment during this period was mainly focused on the expansion of national defence and on state-owned enterprises, which largely transferred private consumption to the army. Second, until the 1970s the economy of Taiwan was confronted with three severe problems: budget deficit, inflationary pressure and confidence problems. The budget deficit was due to the growing fiscal spending on economic reconstruction occurring since World War II. The inflation pressure was related to two oil crises and increasing fiscal spending. The confidence problem was rooted in the political struggle between Taiwan and communist China, emphasised by the cut in diplomatic ties with the US in 1978, which negatively affected the confidence of Taiwan residents. These problems could adversely affect the multiplier effect of expansionary government investment on private consumption.

Between 1970 and 1990, government investment was mainly related to infrastructure construction; the typical example is the Ten Great Economic Constructions. This might be the reason why the crowding-in effect dominates. The rationale is simple: Aschauer (1989a, 1989b, 1990) shows that if infrastructure construction turns down, then total

TABLE 6
Disposable income, final consumption and savings
(NT$)

Year	Disposable Income	Final Consumption	Savings Rate (%)
1964	28591	25381	11.23
1970	44486	40929	8.00
1975	101821	86849	14.87
1980	233112	179687	23.17
1985	320495	246277	23.52
1990	520147	370323	28.80
1994	769755	545987	29.07

(Source: *Report of the Survey of Family Income & Expenditure in Taiwan Area of Republic of China* 1997, Directorate-General of Budget, Accounting and Statistics, Executive Yuan, Taiwan, ROC)

factor productivity turns down slightly later. The point is that infrastructure construction provides a service that minimizes cost: the telecommunication system, for example, improves information technology and enhances production efficiency; and more highways save time spent in traffic.

This section has attempted to explain the outcomes of Taiwan's experience of economic development; however, careful interpretations are required. At any point of time, the explanations are not mutually exclusive, but reflect the synthesis of different forces. Moreover, the standard deviation of regime 1 is 1.359, which is roughly seven times greater than the second regime. This implies that, when the government investment crowds out the private investment, it causes larger investment volatility. Therefore, fiscal expansion should be avoided.

CONCLUSION

In economic theory, the crowding-out effect is largely related to the means that government used to finance an increase in its spending (see Blinder & Solow 1973; Barro 1974; Tobin & Buiter 1980: 204–13). Moreover, Modigliani (1986), Modigliani and Ando (1976) and Stein (1982) discuss demand-management policies and provide econometric evidence. If the multiplier process assumes, as usual, that government sells bonds to finance an increase in its spending; then extra crowding-out comes about in two ways: first, it raises the interest rate. To sell bonds, the government must raise the interest rate to make the bonds attractive. The higher interest rate crowds out all components in aggregate demand. Second, when the bonds mature, interest and principal must be paid to the bondholders. According to the Ricardian equivalence theory, people expect that future taxes will be higher because of this and react by increasing their savings to build up a reserve so that those anticipated higher taxes can be paid without disrupting future consumption levels.

Using the data of Taiwan, the crowding-out hypothesis is appropriately examined by the changes in the exchange rate regime. The study's empirical evidence indicates that fiscal policy is ineffective in a

FIGURE 1
Prime lending rate

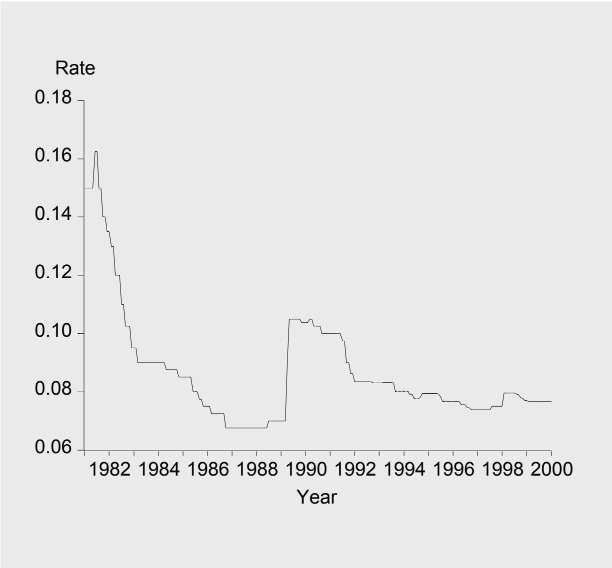


FIGURE 2
One-month deposit rate

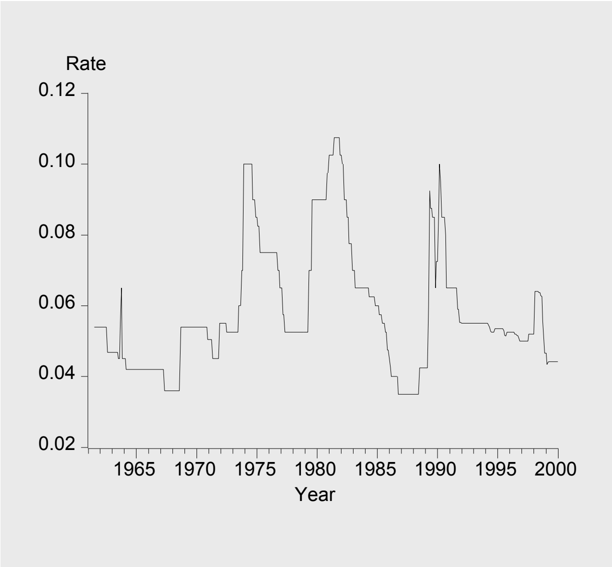


FIGURE 3
One-year deposit rate

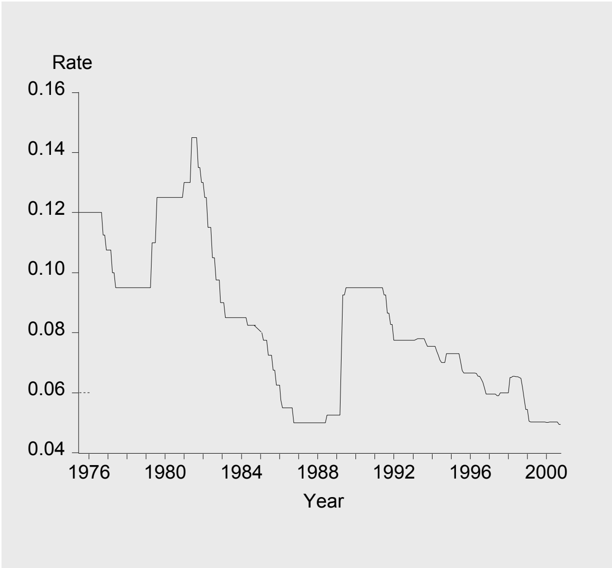


FIGURE 4
Over-night inter-bank rate

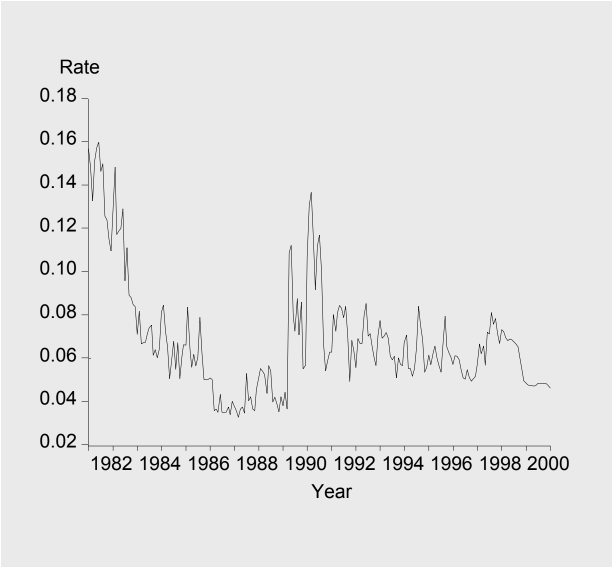


FIGURE 5
Plot of two time series variables

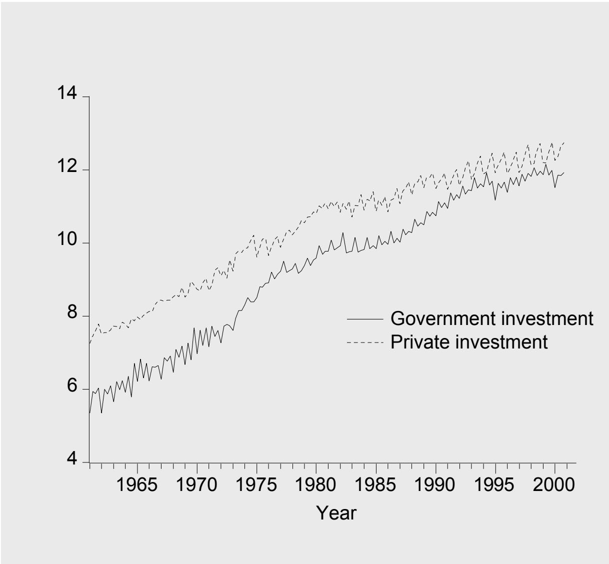


FIGURE 6
Recursive residuals for stability

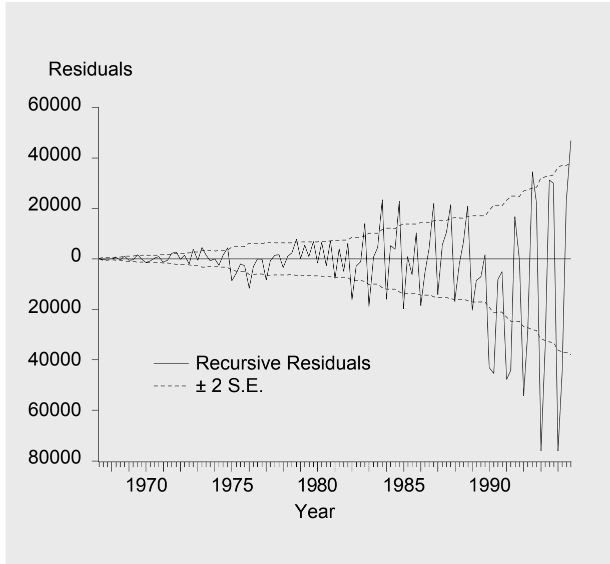


FIGURE 7a
Residuals of full-sample VECM

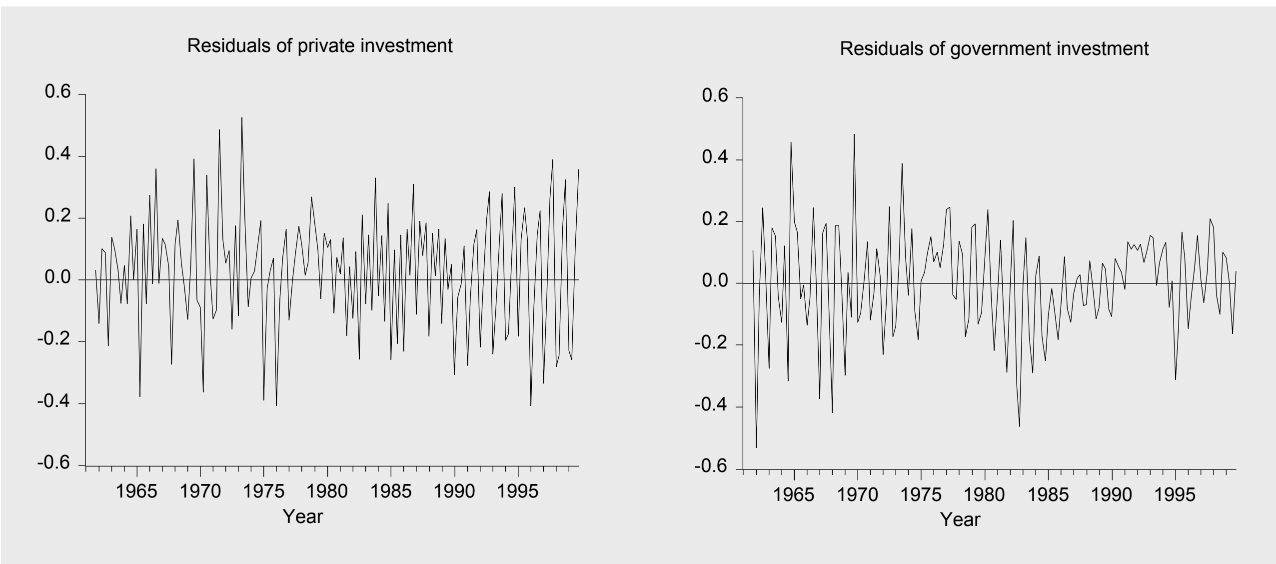


FIGURE 7b
Residuals of full-sample VECM

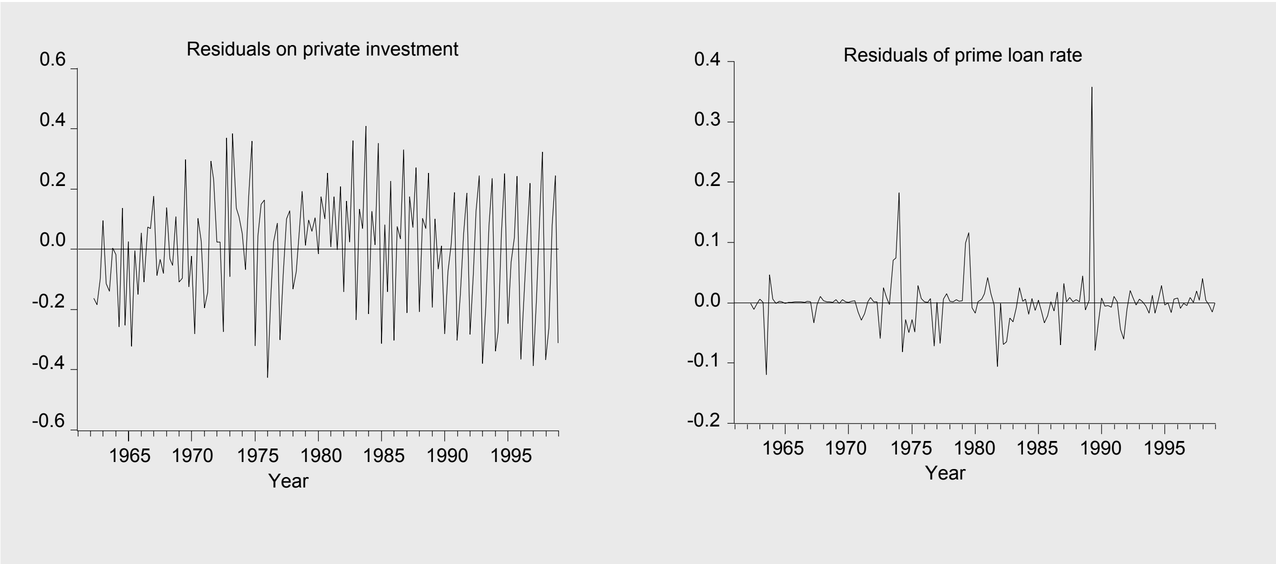


FIGURE 8
Inferred probabilities of the crowding-out
Regime 1(p11)

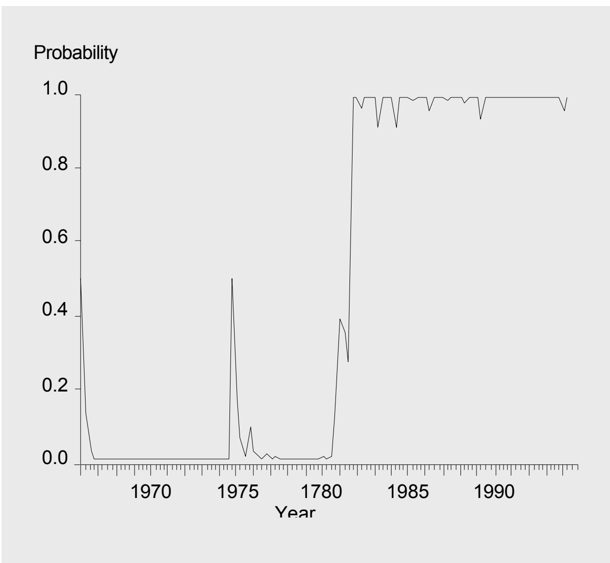


FIGURE 9
The movement of spot exchange rate, NT$/US$

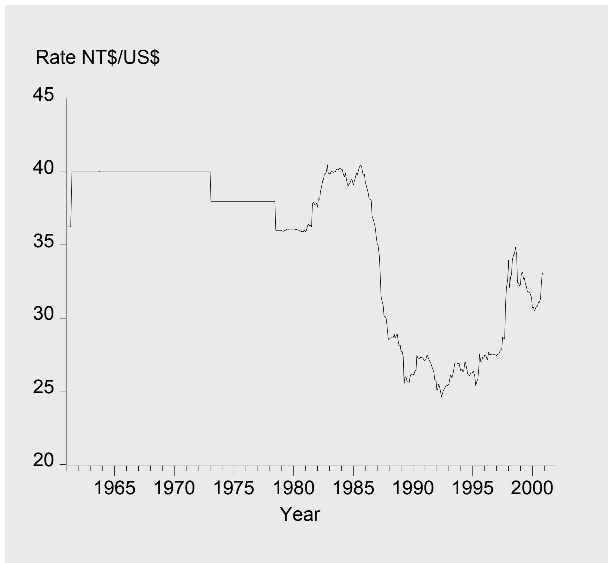
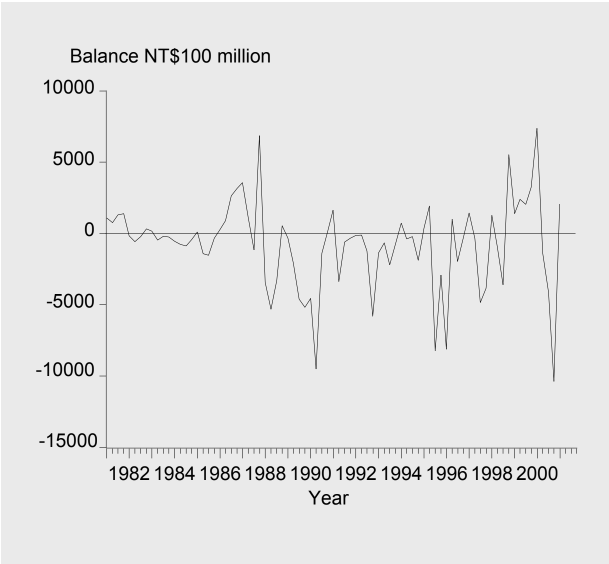


FIGURE 10
Capital account balance
(NT$100 million)



flexible exchange rate regime. The inferred probabilities precisely identify the period that the New Taiwan dollar was allowed to float. Since then, not only has the National Taiwan dollar become flexible, but also the financial market has become more liberalized. In a nutshell, fiscal expansion should be avoided as a stabilization policy.

Moreover, the occurrence of crowding-out has many implications; hence, any interpretations of the reasons behind it must be carefully thought out. For instance, it may imply an inefficient allocation of financial resources; the government's privileged access to scarce resources which increases firms' transaction cost to procure the same services; the private sector's expectation about the effect of fiscal expansion on future investment profitability; and whether the economy is close to full employment. Accordingly, the first conclusion to be drawn is that the ideal policy to minimize the crowding-out effect is to optimize the distribution of government spending for different purposes.

Secondly, the study does not suggest that Taiwan's public sector is inefficient in stimulating the national economy, although it negatively affects most private sectors' investment behaviour, especially the spending on 'investment'. The study shows overall that the crowding-out effects on private investment do not justify the negative effects on the national economy.

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REFERENCES

Ahmed, Shaghill 1986, 'Temporary and permanent government spending in an open economy: some evidences for the United Kingdom', *Journal of Monetary Economics*, 17: 197–224.

Aiyagari, S., Rao, L., Christiano, J. & Eichenbaum, M. 1992, 'The output, employment and interest rates effects of government consumption', *Journal of Monetary Economics*, 30: 73–86.

Amano, R.A. & Wirjanto, T.S. 1997, 'Intratemporal substitution and government spending', *Review of Economics and Statistics*, 79: 605–09.

Aschauer, D.A. 1989a, 'Is public expenditure productive?', *Journal of Monetary Economics,* 23(2): 177–200.

Aschauer, D.A. 1989b, 'Public investment and productivity growth in the Group of Seven', *Journal of Economic Perspective,* 13: 17–25.

Aschauer, D.A. 1990, 'Why is infrastructure important?', in *Is There a Shortfall in Public Capital Investment?,* ed. A.H. Munnell, Conference Series, no. 34, Federal Reserve Bank of Boston, Boston, pp. 21–50.

Banerjee, A.J., Dolado, J., Hendry, D.F & Smith, G.W. 1986, 'Exploring equilibrium relationships in econometrics through statistical models: some Monte Carlo evidence', *Oxford Bulletin of Economics and Statistics*, 48: 253–77.

Barro, R.J. 1974, 'Are government bonds net worth?', *Journal of Political Economy*, 82: 1095–117.

Barro, R.J. 1981, 'Output effects of government purchases', *Journal of Political Economy*, 84: 343–50.

Barro, R.J. 1989, 'The neoclassical approach to fiscal policy', in *Modern Business Cycle Theory*, ed. R.J. Barro, Mass. Harvard University Press, Cambridge.

Baxter, M. & King, R.G. 1993. 'Fiscal policy in general equilibrium', *American Economic Review*, 83: 315–34.

Blinder, A.S. & Solow, R.M. 1973, 'Analytical foundations of fiscal policy', in *The Economics of Public Finance*, A. S. Blinder *et al.*, The Brookings Institution, Washington, pp. 3–115.

Christiano, L. & Eichenbaum, M. 1992, 'Current real-business-cycle theories and aggregate labor-market', *American Economic Review*, 82: 430–50.

Council for Economic Planning and Development of the Executive Yuan of Taiwan, ROC 1961–99, *Taiwan Statistical Data Book*, various issues, Executive Yuan of Taiwan, ROC.

Devereux, M.B., Head, A.C & Lapham, B.J. 1996, 'Monopolistic competition, increasing returns and the effects of government spending', *Journal of Money, Credit and Banking*, 28: 233–54.

Engle, R.F. & Granger, C.W.J. 1987, 'Co-integration and error correction: representation, estimation and testing', *Econometrica*, 55(2): 251–76.

Gustav, R. (ed.) 1992, *Taiwan: From Developing to Mature Economy*, Westview Press, Boulder.

Hall, R.E. 1978, 'Stochastic implications of the life-cycle permanent income hypothesis: theory and evidence', *Journal of Political Economy*, 86: 971–87.

Hamilton, J. 1989, 'A new approach to the economic analysis of nonstationary time series and business cycles', *Econometrica*, 57: 357–384.

Hamilton, J. 1994, *Time Series Analysis*, Princeton University Press, Princeton.

Hamilton, J. 1996, 'Specification testing in Markov-switching time series models', *Journal of Econometrics*, 70: 127–57.

Johansen, S. 1988, 'Statistical analysis of cointegration vectors', *Journal of Economic Dynamics and Control*, 12: 231–54.

Johansen, S. 1990, 'Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models', *Econometrica*, 59: 1551–80.

Karras, G. 1994, 'Government spending and private consumption: some international evidence', *Journal of Money, Credit and Banking*, 26: 9–22.

Kim, C.-J. 1994, 'Dynamic linear models with Markov switching', *Journal of Econometrics*, 60: 1–22.

Kormendi, R.C. 1983, 'Government debt, government spending and private sector behavior', *American Economic Review*, 73: 994–1010.

Kremers, J.J.M., Ericsson, N. & Dolado J. 1992, 'The power of cointegration tests', *Oxford Bulletin of Economics and Statistics*, Aug: 325–48.

Kuo, S.W.Y. 1983, *The Taiwan Economy in Transition*, Westview Press, Boulder, Colorado.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P. & Shin, Y. 1992, 'Testing the null of stationarity against the alternative of a unit root: how sure are we the economic time series have a unit root?', *Journal of Econometrics,* 54: 159–78.

McKibbin, W. & Sachs, J. 1991, *Global Linkage: Macroeconomic Interdependence & Cooperation in the World Economy*, Brookings Institution, Washington, DC.

Modigliani, F. 1986, *The Debate Over Stabilization Policy*, Cambridge University Press, Cambridge.

Modigliani, F. & Ando, A. 1976, 'Impacts of fiscal actions on aggregate income and the monetarist controversy', in *Monetarism*, ed. J.L. Stein, North-Holland, Amsterdam, pp. 17–42.

Norden, S.V. & Vigfusson, R. 1996, Regime-switching models—a guide to the Bank of Canada Gauss procedures, Working paper 96–3, Bank of Canada.

Park, J.Y. 1992, 'Canonical cointegrating regressions', *Econometrica*, 60(1): 119–43.

Phillips, P.C.B. 1987, 'Time series regression with a unit root', *Econometrica*, 55: 227–301.

Phillips, P.C.B. 1991, 'Optimal inference in cointegrated systems', *Econometrica*, 59: 283–306.

Phillips, P.C.B. & Durlauf, S.N. 1986, 'Multiple time series regression with integrated processes', *Review of Economic Studies*, 53: 473–95.

Phillips, P.C.B. & Hansen, B.E. 1990, 'Statistical inference in instrumental variables regression with I(1) processes', *Review of Economic Studies*, 57: 99–125.

Phillips, P.C.B. & Loretan, M. 1991, 'Estimating long-run economic equilibria', *Review of Economic Studies*, 58: 407–36.

Phillips, P.C.B. & Ouliaris, S. 1990, 'Asymptotic properties of residual based tests for cointegration', *Econometrica*, 58: 165–93.

Said, S.E. & Dickey, D.A. 1984, 'Testing for unit roots in autoregressive-moving average models of unknown order', *Biometrica*, 71: 599–607.

Stein, J. L. 1982, *Monetarist, Keynesian and the New Classical Economics*, Basil Blackwell, Oxford.

Tobin, J. & Buiter, W. 1980, 'Fiscal and monetary policies, capital formation and economic activity', reprinted in, *Essay in Economics: Theory and Policy*, James Tobin, The MIT Press, Cambridge, MA, pp. 282–59.

White, H. 1987, 'Specification testing in dynamic models', in *Advances in Econometrics,* vol. II, Fifth World Congress, ed. Truman F. Bewley, Cambridge University Press, Cambridge.

White, H. 1994, *Estimation, Inference and Specification Analysis*, Econometric Society Monograph, Cambridge University Press, Cambridge.

Wickens, M.R. & Breusch, T.S. 1988, 'Dynamic specification, the long run and the estimation of the transformed regression models', *Economic Journal*, 98: 189–205.