THREE ESSAYS ON EMPIRICAL MACROECONOMICS

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Dedication

I wish to thank my wife Elisabetta, my daughter Giulia, my son Kenji, and my parents for their encouragement and emotional support for all the time that I was writing this dissertation.

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I wish to thank Professors Joutz and Trost, and Assistant Professor Sinclair for their patience, guidance, and useful suggestions on my dissertation projects. In particular, I am grateful to Professor Joutz as he always encouraged me to complete this dissertation when I faced challenges. Without his dedicated guidance, I would not be able to accomplish this dissertation. Many thanks to Adjunct Professor Kilpatrick and Assistant Professor Wei for giving me useful comments at my dissertation defense. I also want to thank Prof. Boulier and Ms. Beatty for their various supports.

I am currently at the International Monetary Fund. My dissertation is a pure academic exercise. The views expressed in this dissertation are my own and do not reflect those of the International Monetary Fund.

Abstract of Dissertation

Three Essays on Empirical Macroeconomics

The issues in my research range broadly from inflation in Japan, money demand in Cambodia and Lao P.D.R., to electricity demand in the Philippines. Although these topics vary, the key issues addressed are quite relevant for critical macroeconomic stability issues in these countries.

Essay I (Revising Japanese Inflation Modeling) analyzes the inflation process in Japan. Japanese inflation has remained around zero percent since the late 1990s. This is puzzling, particularly given that none of inflationary factors, such as sustained economic growth and the accommodative monetary policy stance has raised inflationary pressures. I model and perform encompassing tests of the inflation process using a Phillips curve markup framework, separately examining the tradable and non-tradable sectors, and a monetary. My results suggest there has been a structural change in the inflation process.

Essay II (Money Demand, Dollarization, and High Inflation in Cambodia and Lao P.D.R.), investigates local currency demand functions for Cambodia and Lao P.D.R. I employ the Autoregressive Distributed Lag bounds testing approach. Both countries have shared similar historical characteristics, experiencing substantial macroeconomic instability during the 1990s and the high degree of dollarization. My empirical analysis finds a long-run relationship between real money and inflation or a change in exchange rates in Lao P.D.R., but not in Cambodia. Dollarization has been reversed in Lao P.D.R. when policymakers are

committed to macroeconomic stability. The observed relationship is not found in the Cambodian data. In Cambodia, dollarization has continued to increase.

In Essay III (Modeling Electricity Demand in the Philippines), I develop models for the residential and industrial electricity sectors. I use the Johansen cointegration and vector error correction methodology to estimate the long-run and short-run price and income elasticities. The results suggest there is a long-run cointegrating relationship in the residential sector between consumption, income, and stock of electric appliances. In the industrial sector, I find a relation between consumption and GDP on a per worker basis. I compare my model to government's official forecast for electricity demand and find it to be at the upper bound of my forecast range.

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Essay I: Revising Japanese Inflation Modeling

I. Introduction

Japanese inflation has remained around zero since the late 1990s. Neither sustained economic growth nor high oil and commodity prices seem to have raised inflationary pressures evidently. The low inflation in Japan is puzzling particularly given that the Bank of Japan has conducted fairly expansionary monetary policy over the last decade, including a zero interest rate policy (1999-2000) and thereafter a quantitative easing policy (2000-2006)[1].

Although not as excessive as Japan, low and stable inflation has been a global phenomenon. Over the past ten years, global inflation has dropped from around 30 percent to low single digits. In recent years, a surge in global oil and commodity prices has raised concerns about the reemergence of inflationary pressures among central bankers, but inflation in many advanced countries has stayed at relatively low levels by historical standards. Accordingly, there has been an increasing volume of literature analyzing whether there has been a change in the inflation process, and if so, what has been behind.

My research attempts to model the inflation process in Japan, with particular focus on analyzing factors behind low inflation. My analysis covers the period of 1970Q1-2007Q1 using quarterly data. There could be a number of potential reasons to explain low inflation,

[1] One of the objectives of this quantitative easing policy was to bring down the whole interest rate term structure, with the short-ends kept at the zero percent level.

but instead of attempting ambitiously to find out which factor most fundamentally matters, this essay takes a practical approach and defines this question narrowly. It analyzes the inflation process in the context of a Phillips curve markup type model, including the separate analysis on tradable sector inflation and non-tradable sector inflation, and a monetary type (real money gap) model, and then investigates whether there has been a structural break in the relationship between inflation and its determinants.

The structure of this paper is as follows. Section II provides a brief literature review. Section III presents Phillips-curve markup type inflation modeling, showing that (i) the inflation process in the tradable and non-tradable sectors can be better modeled separately, and (ii) the evidence of a structural change in markup is found. Section IV discusses that a monetary type inflation model can also explain inflation in Japan well. Section V compares a Phillips curve markup type model and a monetary type model and shows that the former model can dominate the latter model. Section VI discusses policy implications based on the empirical models, and Section VII concludes.

II. Related literature

In contrast to the research in this area 20-30 years ago, where research efforts were motivated by high inflation in the aftermath of two major oil shocks in the 1970s and the early 1980s, the recent literature attempts to tackle quite an opposite issue: why inflation has declined and remained broadly stable generally in advanced countries? This said, interest in research in this line apparently somewhat receded in most recent years, reflecting a concern about a revival in inflation arising from higher oil and commodity prices.

However, with the recent outbreak of a major financial crisis in the U.S. and globally and amounting concerns about a recession to come, I expect that researchers' focus would again turn to disinflation or even deflation.

A group of researchers attempted to attribute the lower and stable inflation to sound monetary policy, including the adoption of inflation targeting as well as greater independence of central banks. For example, see Bernanke et at (1999), Johnson (2002), Mishkin and Schmidt-Hebbel (2002), and International Monetary Fund (2003).

The other group emphasized that the lower inflation had been due to increased productivity. Ball and Moffitt (2001) and De Long (2002) argued that unexpected productivity growth was one of the main factors for disinflation in the United States. However, for the other countries, the evidence was weak: Gust and Marquez (2002) studied a panel of OECD countries and found that only two out of the 20 OECD countries in the sample experienced an acceleration in labor productivity in the 1990s.

Some others pointed to the labor market reform, typically focusing on the possibility of a decline in the Non-Accelerating Inflation Rate of Unemployment (NAIRU). For example, Ball and Mankiw (2002) argued that structural changes in labor markets, due to changes in demographics or changes in government policies regarding disability insurance, had reduced the NAIRU in the United States. See also, among others, Staiger et al (2001), Fabiani and Mestre (2001), and Greenslade et al (2003).

Finally, the possible effect of globalization on inflation has also been analyzed. Rogoff (2003) stressed the impact of globalization on disinflation, one of the most influential studies in linking inflation and globalization. Others included Nickell (2005) emphasizing the deflationary impact arising from the production switch to low cost countries like China; and Ihrig et al (2007) evaluating the hypothesis that globalization had increased the role of international factors and decreased the role of domestic factors in the inflation process in industrial countries.

Studies on Japan's disinflation experience included Tanaka and Kimura (1998), Sekine (2001), Kamada and Hirakata (2002), and Shibamoto and Miyao (2008). Tanaka and Kimura (1998) demonstrated that a Phillips curve markup type model explained the inflation process in Japan reasonably well. Sekine (2001) compared the inflation forecasting performance of alternative models, including Phillips curve, monetary, and exchange rate type models. Kamada and Hiraoka (2002) incorporated a structural change (import penetration) into inflation modeling, demonstrating that the structural change mattered for a change in the inflation process in Japan. In the most recent study, Shibamoto and Miyao (2008) attempted to explore possible explanations for stable Japanese inflation based on an aggregate demand and supply framework using VAR.

III. Phillips-curve Markup Type Inflation Modeling

A. Modeling Strategy

General issues

In line with Franz and Gordon (1993), Cockerell and Russel (1995), De Brouwer and Ericsson (1998), earlier literature on a Phillips curve markup type model for Japan (such as Tanaka and Kimura, 1998, and Sekine, 2001) assumed that the long-run price level (CPI) is determined by markup over unit labor costs (*ULC*):

(1) $CPI = \mu(ULC^{\beta})$.

Tanaka and Kimura (1998) estimated equation (1) as a two-variable cointegrating VAR system, while Sekine (2001) added input prices (the whole price index) into a long-run relationship, $CPI = \mu(ULC^{\beta 1})(IP^{\beta 2})$, which he estimated as a three-variable cointegrating VAR system.

However, these earlier models may not prove sufficiently robust to changing economic environment and structural shocks over the last decade. In fact, updating these earlier models with recent data suggests parameter instability or even weaker evidence of such a long-run relationship, which I will first empirically demonstrate in this section. Such unfavorable results could reflect its simplicity in modeling strategy—although simplicity *per se* can be regarded as success provided that an estimated model is robust. I argue that the assumption of constant markup $\mu - 1$ over total costs does not accord with structural shocks and thus should be reviewed. It would be more natural to assume that markup may

fluctuate due to firms' pricing behavior, which would in turn reflect overall aggregate demand conditions, their monopolistic competition power, etc.

Against this backdrop, this section attempts alternative approaches to Japanese inflation modeling by explicitly embedding possible variables that could be considered to affect firms' markup behavior. Those variables (hereafter, called markup factors) are chosen *a priori* based on economic theory and recent studies on disinflation (as summarized in the related literature), taking account of broadly-shared perceptions about factors containing inflation pressures in Japan over the last decade. However, in empirics, estimating a large unrestricted VAR model, which incorporates all possible variables that matter for inflation, is quite challenging to handle. It is often the case that in the context of the analysis of a multivariate cointegrating model, the difficulty of interpreting the cointegration space grows as more variables are added to VAR.

As one solution to this practical difficulty, I propose a sector approach. Under the hypothesis that the data generating process between the tradable and non-tradable sectors differs, I argue that estimating each sector separately could help reduce the size of VAR. This approach would also lead to a more accurate model specification than the one estimated through the traditional aggregation approach (i.e., estimating total inflation), because it takes better account of individual dynamic properties (Theil, 1954).[2]

[2] However, Grunfeld and Griliches (1960) argued that the sector approach would not be necessarily superior to the aggregation approach, particularly if there was doubt about how correctly a model was specified. Hubrich (2005) noted that which approach outweighs the other would largely be an empirical question.

The following variables are chosen *a priori* as possible markup factors. In the tradable sector, the focus is on factors affecting competition with foreign countries, following Rogoff (2003) who stressed the impact of globalization on disinflation. In this context, among several possible measurements of competitiveness, I look at an openness indicator (the sum of exports and imports as a percent of GDP) and real exchange rates. After Plaza and Louvre Accords which led to sharp appreciations of Japanese yen during the late 1980s, Japanese manufactures incorporated themselves aggressively into the Asian international production network by shifting their production lines into low labor cost Asian countries and importing final goods from them. As a result, an increase in cheap imports directly affected consumer prices, and also indirectly through raising competition pressures in the tradable sector. Thus, an increase in openness should have affected prices in Japan. The effect of the level of real exchange rates on competitiveness is widely acknowledged in international economics literature, and thus no additional clarification should be warranted.

In the non-tradable sector, factors relevant to labor market conditions and asset markets are examined. On the first factor, given that workers in Japan were long accustomed to lifetime employment prospects during the post war periods, they have become extremely pessimistic about their job security in the face of rising unemployment rates starting in the middle of the 1990s. The increased unemployment rate likely reduced people's willingness to consume, resulting in a decline in firms' markup. On the second factor, a hypothesis is

that a sharp drop in asset prices starting in the early 1990s should have dampened consumers' confidence, contributing to firms' lower markup.

Econometric issues

I interpret a Phillips curve-type inflation model, like equation (1) and (5) (in subsequent Section III. D.), as describing an equilibrium relationship among the variables of interest (e.g., *CPI*, *ULC*, *IP*, and *Markup Factor* to be discussed later). To put it more precisely in empirical context, the model shows a steady state relationship among these variables, given that these variables are non-stationary processes. In other words, if these variables are linked to form an equilibrium relationship in the long-run, they will move closely together over time.

This interpretation is consistent with the objectives of my research because my main interest lies in the determinants and behavior of inflation and its interactions with other relevant variables (e.g. *ULC*) in the long-run. My research attempts to test whether any equilibrium Phillips curve type relationship exists. If so, whether it is stable, and if not, whether a breakdown of the relationship can be accounted for by a structural change or missing variables or information that was not well captured in the previous econometric analysis by other researchers.

Accordingly, the cointegration approach, suggested by Engel and Granger (1987) and Johansen (1988 and 1991), would be most suited for my research objectives. Between Engel and Granger's single equation approach and Johansen's multivariate system

approach, I choose the latter approach. This is primarily because Johansen's VAR approach allows me to build a model without imposing any assumptions about endogeneity issues: whether changes in any of the right-hand-side variables in equation (1) or (5) cause changes in the left-hand-side variable *CPI*. Besides, I will employ a general-to-specific approach: I start with an explicit stochastic system formulation of all *relevant* data and then reduce the general dynamic model by imposing testable restrictions on the parameters. Here, the relevant data are chosen based on my theoretical interest and hypothesis—e.g., whether Phillips curve-type model explains Japanese inflation well and whether alternative variable(s) affecting firms' markup behavior should be embedded in the model. Then I will attempt to answer these questions based on a statistical analysis of the VAR model. Finally, I am also interested in the short-run dynamics of the variables under consideration, especially because the steady state may rarely be observed. While there is adjustment back to steady state it is continuallydisturbed by new shocks. In sum, I will attempt to develop an empirical model where the long-run determinants of inflation are accounted for by a cointegration relationship, and the short run determinants by a vector error correction model or Autoregressive Distributed Lag (ARDL) model. The application of these techniques to inflation modeling has been fully developed in Brouwer and Ericsson (1998) for the case of Australia and Tanaka and Kitamura (1998) and Sekine (2001) for the case of Japan.

B. Key Data and Integration Property

All data are quarterly and available for 1970Q1-2007Q1. However, observations for the first couple of quarters may be left aside for lags. Main variables of interest in this section are inflation, unit labor cost, input prices, markup factors, and a GDP gap indicator.

Inflation (*dcpi*) is defined in this paper as growth of the consumer price index excluding fresh foods and petroleum products (*cpi*).[3] Fresh foods and petroleum products are excluded from total CPI because the former is subject to large temporal volatility, while the exclusion of the later is to investigate possible impacts of energy prices on the inflation process. Growth of CPI for goods (excluding fresh foods and petroleum products) is used as proxy for inflation in the tradable sector (*cpit*), while growth of CPI for services is used as proxy for inflation in the non-tradable sector (*cpint*).

Unit labor cost (*ulc*) is defined as the log difference of compensation of employees (*com*) and GDP output (*y*), $ulc_t = com_t - y_t$ [4].

Input prices include the overall import price index (*ip*) as well as the import price index for energy (*ipe*). All of these data are from the whole price index[5].

[3] A small letter indicates the logarithm, unless otherwise indicated, and *d* indicates the first difference.

[4] Source: Department of National Accounts, ESRI, Cabinet Office, http://www.esri.cao.go.jp/en/sna/menu.html and Statistics Bureau and Statistics Center, Ministry of Public Management, Home Affairs, Posts and Telecommunications, http://www.stat.go.jp/english/data/cpi/index.htm.

Openness (*openness*) is defined as the ratio of exports and import of goods and services as a percent of GDP. Data for exports and imports are from Japanese national accounts. Real exchange rates (*reex*) are real effective exchange rate indices calculated by the Bank of Japan.[6] Unemployment rate (*un*) is defined as unemployed (persons who did not work at all during the survey week, but were currently available for work and were actively seeking a job or were waiting for the results of the past job-seeking activity) divided by labor force (in percent)[7]. As asset prices, Nikkei 225 stock exchange indices (*stock price*) and the real estate stock price indices (*estate stock price*)—the average of seven major real estate companies weighted by their market value as proxy for the housing price—are used.

Finally, GDP gap (*gap*) is estimated as the deviation of the actual GDP from its potential. Potential GDP (y*) is from a production-based approach (OECD estimates).[8]

Figure I.1 plots the levels of *CPI*, *CPIT*, and *CPINT*. All these measures of prices had increased until the early 1990s, but have since declined or remained flat. Table I.1 shows that the statistical property of these variables (mean, standard deviation, and skewness) is clearly different. In terms of annual percentage increase, all these variables followed

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[5] Units: 1995=100. Source: Research and Statistics Department, The Bank of Japan, http://www.boj.or.jp/en/siryo/siryo_f.htm.

[6] Units: 1996=100. Source: Research and Statistics Department, The Bank of Japan, http://www.boj.or.jp/en/theme/research/stat/market/forex/jikko/index.htm.

[7] Source: Labor force survey, Statistics Bureau and Statistics Center, Ministry of Public Management, Home Affairs, Posts and Telecommunications, http://www.stat.go.jp/english/data/index.htm.

[8] Source: OECD.Stat, http://stats.oecd.org/wbos/Index.aspx?usercontext=sourceoecd.

downward trends after the early 1970s, and are currently hovering around zero inflation rates (Figure I.2).

Figures I.3 and 4 plot the level and growth rate of ULC. The level of ULC closely follows the trends of CPI indicators until the mid-1990s. However, it has since faced a shaper decline than CPI indicators, reflecting a decline in compensation of employee with steady growth of productivity.

Figures I.5 and 6 show levels and growth of input prices, *IP* and *IPE*. Of interest is how recent increases in energy prices have affected inflation in Japan. The following two facts are pointed out immediately from these figures. First, import prices (both levels and annual growth) are highly volatile, compared to CPI inflation. Second, the correlations between CPI indicators and import prices look generally weak. For example, import prices rose sharply in 1981 (the second oil shock) and dropped in 1987 (a sharp yen appreciation after the Louvre Accord), but CPI indicators did not swing so widely. Furthermore, import prices have risen markedly since 2005, but CPI indicators have remained virtually flat.

Panel Figure I.7 summarizes various markup factors and GDP gap. The openness indicator has risen sharply since the early 1990s, reflecting increased integration of Japanese economy into the neighboring Asian countries. Besides, the real effective exchange rate (*reer*) climbed to elevated levels through the 1990s, which should have reduced the competitiveness of the tradable sector in Japan. Unemployment rates increased steadily to unprecedented levels in the 1990s, which dramatically changed workers' perceptions about

job security, contributing to a reduction in spending (and thereby, lower firms' markup). A decline in wealth—both overall stock price index and estate stock price index fell steadily during the 1990s—should have also contributed to weaker demand and lower firms' markup. Finally, GDP gap presents relatively clear cyclical fluctuations, with its levels mostly falling below zero during the recent low inflation periods of 1998-2006.

As most of these variables are seemingly non-stationary, their orders of the integration should be tested. Tables I.2 and 3 summarize the results of Augmented Dickey Fuller (ADF) unit root test and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) stationarity test. For each variable, I start with a very general model specification including a trend and eight lags, and then simplify it by dropping irrelevant tend or lag variables.

ADF tests show that the levels of all CPI and unit labor cost variables are stationary. This result is surprising and not consistent with the earlier studies (for example, Tanaka and Kimura, 1997, and Sekine, 2001).[9] Note that estimated coefficients on the lagged value, γ-*1*, are quite close to zero, questioning the robustness of the test results. Thus, the alternative approach, KPSS, is employed. The results show that the null hypotheses of stationarity can be rejected for all these variables. For the first difference of these variables, KPSS tests indicate that all these variables are clearly stationary—although for *cpi*, the null hypothesis of stationarity of its first difference cannot be rejected only at the 5 percent significance

[9] This is mainly due to difference in sample periods. If ADF test is employed for a shorter sample period starting 1980 (as in Tanaka and Kimura) or ending 1998 (as in Sekine), the null hypothesis of the unit root cannot be rejected.

level. Against this background, in the following analysis, these CPI and unit labor costs variables are assumed to be I(1).

ADF tests also show that all input prices (*ip* and *ipe*) and markup factors (*openness, reer, un, stock price,* and *estate stock price*) are I(1). KPSS tests confirm non-stationarity of these variables, except for *ip*, but given the small estimated value of γ*-1* in the ADF test, I assume that *ip* is I(1). For GDP gap, ADF and KPSS tests lead to conflicting results with the former test supporting its nonstationarity and the latter its stationarity. However, given that the estimated value of γ*-1* is relatively large at 0.08, I assume that *gap* is I(1).

With integration property of the data clarified, I move to the estimation of a long-run Phillips curve markup type model. The next sub-section C re-estimates a traditional long-run Phillips curve markup model studied in the earlier literature, and sub-section D presents an alternative sectoral approach taking account of a different data generation process for the tradable and non-tradable sectors and explicitly recognizing markup factors.

C. Traditional Phillips Curve Markup Type Inflation Modeling

Model

Adding input prices (*IP*), equation (1) can be rewritten as:

$$(2)\quad CPI = \mu(ULC^{\beta 1})(IP^{\beta 2}).$$

For input prices, import price indices for overall import (*IP*) and energy (*IPE*) will be considered. In the log linear form, equation (2) is:

$$(3)\ cpi = \ln \mu + \beta_1 \cdot ulc + \beta_2 \cdot ip\ ,\ \beta_1, \beta_2 > 0\ ,$$

where lowercase letters show the logarithms of the variables.

Estimation

The estimation of model (3) consists of the three steps: (i) selecting the lag order of VAR; (ii) estimating long run equations by the Johansen cointegration test; and (iii) testing parameter restrictions such as weakly exogeneity, if warranted.

The first step is to determine the lag order of VAR. The following VAR is estimated:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + B\ x_t + \varepsilon_t$$

where y_t = the $k \times 1$ vector of non stationary I(1) variables,

x_t = the $l \times 1$ vector of deterministic (exogenous) variables,

A_i = the $k \times k$ matrix of coefficients,

B = the $k \times l$ matrix of coefficients,

ε_t = the $k \times 1$ vector of error terms.

Exogenous variables include constant, trend, first oil shock dummies (*oil74q2*, 1974Q2=1), consumption tax dummies (*tax89*, 1989Q2=1 and *tax97*, 1997Q2=1), high growth period dummies (*highgrowth*, 1971Q1-2003Q3=1), and zero interest rate period dummies (*zeroint*, 1999Q1-2006Q1=1)[10]. The consumption tax dummy variables take account of (i) the introduction of 3 percent consumption tax in April 1989 and (ii) a further increase in the tax rate from 3 to 5 percent in April 1997.

Three VAR models are estimated: a model with two endogenous variables (*cpi* and *ulc*), a model with three endogenous variables (*cpi*, *ulc*, and *ip*), and a model with (*cpi*, *ulc*, and *ipe*). The maximum lag order to begin with is eight. Table I.4 summarizes the statistics of the Akaike Information Criterion (AIC), Hannan-Quinn Criterion (HQ) and Schwarz Criterion (SC) from lag one to eight for each of the three VAR models. For all cases, AIC tends to choose the largest lag order, while SC the shortest lag order. As known, the SC is superior in large sample properties and asymptotically consistent, whereas the AIC is biased toward selecting an over parameterized model. However, it is not clear whether the results here satisfy large sample properties, and therefore, diagnostic tests are run for all these possible cases to judge an appropriate lag order.

Table I.5 reports diagnostic statistics for the three VAR models. For each VAR model, two models with a short lag order and a long lag order are tested (the tests include error autocorrelation, normality, and heteroskedasticity). The test results are not generally

[10] The second oil shock dummy (1980Q2=1) and Gulf war dummy (1990Q4=1) were also considered. However, these dummies were found to be insignificant, and therefore

satisfactory: all models have error serial autocorrelation problems, regardless of lag length. However, if the sample period prior to 1980Q1 is omitted, almost all models have satisfactory properties for the residuals as problems associated with autocorrelation, normality, and homoskedasticity largely disappear (although not fully). This implies that the unsatisfactory results of diagnostic statistics for the full sample periods would be related to volatile movements of the endogenous variables during the 1970s, in part reflecting oil shocks.

Accordingly, a compelling argument can be made that analysis should be limited to shorter sample periods. However, the observations of the 1970s contain important information about the experience of oil shocks, an issue that Japan is exactly facing in recent years. Furthermore, a larger sample size is considered generally suitable for estimating a multivariable cointegration model. Accordingly, the following analysis primary focuses on the full sample period, although a shorter sample period is also considered, as needed, for checking the robustness of estimated models, given concerns about VAR residual problems.

The results of Johansen cointegration analysis are reported in Table I.6. The cointegration relations are assumed to have a linear trend, and the trend is highly significant with a negative coefficient. The inclusion of the trend term partly reflects a fact that labor income ratio follows a negative trend. For all three models, the trace tests indicate the existence of

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dropped.

at most 1 cointegrating equation at the 1 percent significance level. However, both overall import prices, *ip*, and energy import prices, *ipe* (models 1 and 2 in Table I.6) have a positive coefficient sign, contradicting the theoretical prediction (equation (2)). Thus, a model that is consistent with the theoretical prediction is only the two variable VAR with (*cpi* and *ulc*), model 3.

$$(4)\quad cpi = \underset{(0.043^{**})}{0.548}\, ulc + \underset{(0.000^{**})}{0.002}\, t\,,$$

where ( ) indicates standard errors. The corresponding α vector is:

$$\begin{matrix} cpi \\ ulc \end{matrix} \begin{pmatrix} -0.10^{**} \\ 0.10^{**} \end{pmatrix}.$$

where ** indicates significance at the 1 percent level. The α coefficient on *cpi* is negative, while that on *ulc* is positive, suggesting an error correction mechanism at work, and both these estimated coefficients are highly significant.

The obtained model (4) is similar to the one reported by Tanaka and Kimura (1998), although their estimated β coefficient on *ulc* is much larger at 0.871, likely reflecting a different sample period under investigation (Tanaka and Kimura analyzed a sample period for 1981Q1-1998Q1). In fact, if the same model is estimated for sample period 1981Q1-1998Q1, the β coefficient on *ulc* becomes 0.820, closer to Tanaka and Kimura's result.[11]

[11] However, in contrast to Tanaka and Kimura's reporting, trend was found insignificant, and thus excluded from the regression.

Sekine (2001) reported a much larger β coefficient on *ulc* (0.90), using the sample period of 1972Q1-1999Q1 and including the whole price index as an input price.[12]

Model (4) is appealing due to its parsimoniousness, but its robustness is in question. To verify this, the same cointegrated VAR is estimated recursively, starting 1985Q1. Figure I.8 shows evidence of parameter instability during the 1990s, indicating the possibility of a structural break.

D. Sectoral Phillips Curve Type Inflation Modeling

Model

Two modifications to equation (3) are attempted: (i) taking an explicit account of markup factors and (ii) estimating it separately for the tradable and non-tradable sectors. Layard, Nickell, and Jackman (1991) argued that the markup tended to rise with the level of activity. More generally, the markup may fluctuate reflecting a factor that affects firms' pricing behavior. Hence, equation (3) can be modified as follows:

(5) $CPI = \mu(ULC^{\beta 1})(IP^{\beta 2})(Markup\ Factor^{\beta 3})$

where *Markup Factor* is the factor that affects the markup, including the intensity of competition with foreign countries, labor market conditions, and asset markets, as discussed previously. Now, a long-run inflation model for the tradable and non-tradable sectors (*cpit* and *cpint*, respectively), corresponding to equation (5), can be expressed as:

(6) $cpit = \ln \mu + \beta_1 \cdot ulc + \beta_2 \cdot ip \ + \beta_3 \cdot markup\ factor^T$, $\beta_1, \beta_2 > 0, \beta_3 < 0$,

[12] I have also tested the whole price index as an input price, but could not find a meaningful

(7) $cp\,\text{int} = \ln \mu + \beta_1 \cdot ulc + \beta_2 \cdot ip \ + \beta_3 \cdot markup\ factor^{NT}$,

$\beta_1, \beta_2 > 0, \beta_3(un) < 0, \beta_3(stock\ price; estate\ stock\ price) > 0$,

For tradable sector $markup\ factor^{T}$, the ratio of exports and imports of goods and services as a percent of GDP (*openness*) and the level of real effective exchange rates (*reex*) are considered, while for non-tradable sector $markup\ factor^{NT}$, the level of unemployment (*un*), nominal stock prices (*stock price*), and nominal real estate stock prices (*estate stock price*) are examined.

Finally, the overall price level can be derived as:

(8) $cpi = \rho \cdot cpit + (1-\rho) \cdot cp\,\text{int}$, $\rho > 0$,

where ρ is the weight of the tradable sector in the economy.

Estimation

Tradable sector

I follow the same process as earlier: first VAR lag selection, then Johansen cointegration tests, and finally parameter restriction tests. The final results of Johansen cointegration tests are reported in Table I.7.[13] The important hypothesis here is whether markup factors, *openness* and *reer*, matter in forming part of cointegrating relationships with *cpit* and *ulc*. Main results can be summarized as following.

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cointegration relationship.

[13] The results of VAR lag selection are reported in Appendix Table I.1.

First, the unit labor cost, *ulc*, remains an important determinant to the long-run relation. Its β coefficients for all specifications are similar to the one obtained for the traditional Phillips curve markup type model.

Second, similar to the case of the traditional model, β coefficients on input pries, *ip* or *ipe*, show positive signs, contradicting the theoretical prediction (models 2 and 3 in Table I.7). This strengthens the earlier evidence from the traditional model that neither import prices nor imported energy prices would mater for the long-run inflation process, even for the tradable sector alone.

Finally, a markup factor, *openness*, is highly significant with a positive coefficient, suggesting that greater openness would reduce the level of *cpit* in the long run (model 4). On the contrary, the real effective exchange rate, *reer*, has a negative coefficient (model 5), implying that a higher level of *reer* (i.e., a more appreciated level of *reer*) would push up the level of *cpit*. This is not immediately interpretable. Given that the real exchange rate represents a relative price of the non-tradable sector compared to the tradable sector, an appreciation of the real exchange rate should have led to a decline in tradable sector prices.[14]

[14] I have also tested whether markup factors for the non-tradable sector, *un*, *stock price*, or *estate stock price*, could be entered into a long-run cointegrating relationship in the tradable sector. However, I could not find a meaningful cointegrating relationship for those variables.

As a conclusion, model 4 that contains a long-run cointegrating relationship among *cpit*, *ulc*, and *openness* satisfies all conditions, and therefore is selected as a tradable sector long run inflation model.

$$(9)\ cpit = \underset{(0.032^{**})}{0.533}\, ulc - \underset{(0.049^{**})}{0.329}\, openness + \underset{(0.000^{**})}{0.003}\, t\ ,$$

The corresponding α vector is:

$$\begin{matrix} cpit \\ ulc \\ openness \end{matrix} \begin{matrix} ... \\ ... \\ ... \end{matrix} \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \end{pmatrix} = \begin{pmatrix} -0.163^{**} \\ 0.127^{*} \\ -0.433^{**} \end{pmatrix}.$$

The α coefficients on *cpi* and *openness* are negative, while that on *ulc* is positive—consistent with error correction mechanisms— and both are highly significant. Weak exogenousty tests for the *ulc* or *openness* equation— i.e., $\alpha_{21} = 0$ and $\alpha_{31} = 0$, respectively—does not hold (models 4a and 4b), supporting the variability of this three-variable cointegrating VAR model.[15]

I test the stability of the parameters, equation (9) is estimated recursively, starting 1985Q1. Besides, to show how this three variable cointegrating VAR model performs well in terms of parameter stability, a traditional two-variable cointegrating VAR model (model 1 in Table I.7) is also recursively estimated. Figures I.9a and 9b show a remarkable contrast between the two models. While estimated β coefficients on *ulc* in the latter model that excludes *openness* are highly instable, notably around 1998 (Figure I.9a), β coefficients on

[15] I have run a VAR regression for a shorter sample period, 1981Q1-2007Q1 and have obtained a broadly similar result.

ulc in equation (9) are remarkably stable (Figure I.9b). This analysis implies that the omission of *openness* indicator led to the parameter instability. This is consistent with my prior hypothesis that an increase in openness contributed to driving the markup in the tradable sector price down, thereby contributing to lowering prices.

Non-tradable sector

For the non-tradable sector, *un*, *stock price*, or *estate stock price*, is included in cointegrating VAR as candidate markup factors. Table I.8 reports the main results.[16] The main findings include, first, β coefficients on *ulc* are negative and are highly significant for all models, although its size varies. In all models, adjustment coefficients on *ulc* equations have a theoretical consistent positive sign but are not significant. This suggests that for the non-tradable sector, *ulc*, is weakly endogenous. Second, again, β coefficients on input pries, *ip* or *ipe*, are positive, contradicting the theoretical prediction (models 2 and 3 in Table I.8). Thus, neither import prices nor imported energy prices would affect the long-run inflation process in the non-tradable sector at all. And finally, two markup factors, *un* and *stock price*, have positive and negative coefficients, respectively, consistent with the prediction (models 4 and 5). Also, these two coefficients are significant at the 5 percent level. However, *estate stock price*, is found to be not significant at the 5 percent level (model 6).[17]

[16] The results of VAR lag selection are reported in Appendix Table I.2.

[17] I have also tested whether markup factors for the tradable sector, *openness* and *reer* could be entered into a long-run cointegrating relationship in the tradable sector. However, I could not find a meaningful cointegrating relationship for those variables.

The above results suggest that newly introduced markup factors, *un* and *stock price*, comprise a long-run relationship together with *cpint* and *ulc*. This means that markup in the non-tradable sector would decline as unemployment rises and stock prices fall. Note that α adjustment coefficients for *ulc* and *un* equations in model 4 and those for *ulc* and *stock price* equations in model 5 are not statistically significant. Likelihood ratio tests for weak exogenousty confirm that the null hypothesis of the exclusion of these parameters cannot be rejected with high probability (restricted cointegration models 4 and 5 in Table I.8). Accordingly, the obtained models are:

$$(10)\ cpint = \underset{(0.036^{**})}{0.710}\, ulc - \underset{(0.005^{*})}{0.012}\, un + \underset{(0.000^{**})}{0.003}\, t\ ,$$

with the corresponding α vector,

$$\begin{matrix} cpi \\ ulc \\ un \end{matrix} \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \end{pmatrix} = \begin{pmatrix} -0.150^{**} \\ 0.000 \\ 0.000 \end{pmatrix},$$ [18] and

$$(11)\ cpint = \underset{(0.048^{**})}{0.561}\, ulc + \underset{(0.009^{**})}{0.026}\, stock\ price + \underset{(0.000^{**})}{0.003}\, t\ ,$$

with the corresponding α vector,

$$\begin{matrix} cpi \\ ulc \\ stock\ price \end{matrix} \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \end{pmatrix} = \begin{pmatrix} -0.103^{**} \\ 0.000 \\ 0.000 \end{pmatrix}.$$ [19]

[18] This cointegration holds for a shorter sample period, 1981Q1-2007Q1 but estimated β coefficient on *ulc* is found to be much larger (around one).

[19] Unfortunately, this cointegration did not hold for a shorter sample period, 1981Q1-2007Q1. Thus, this specification is not robust to the selection of sample periods.

Recursive estimates also confirm that the estimated β coefficients on *ulc* for the above two models are broadly stable (Figures I.10b and 10c). This is a noteworthy contrast to a simple two-variable cointegrating VAR model with *cpint* and *ulc* (model 1 in Table I.8), showing some instability in the early 1990s (Figure I.10a). This result implies that a shift in unemployment trend and stock prices caused a structural change in the long-run price of the non-tradable sector.

Overall price level equation

Finally, substituting the estimated tradable and non-tradable sector price models, $\overline{cpit}$ (equation 9) and $\overline{cpint}$ (equation 10 or 11), into equation (8), I get the following long-run models for the overall price levels:

(12) $cpi = \rho \cdot \overline{cpit} + (1-\rho) \cdot \overline{cpint}$

A weight of the tradable sector, ρ, is 0.4, from the CPI statistics.[20] Then, equation (12) can be rewritten as:

Combination of equations (9) and (10)

(13) $cpi = 0.4 \cdot \{0.533\,ulc - 0.329\,openness + 0.003\,t\} + 0.6 \cdot \{0.710\,ulc - 0.012\,un + 0.003\,t\}$

$$= 0.639 ulc - 0.132 openness - 0.007 un + 0.003 t$$

Combination of equations (9) and (11)

[20] The weights of the CPI statistics are calculated on the basis of average expenditures per household derived from the Family Income and Expenditure Survey. Over the last decade, the weights were revised in 2005, 2000, and 1995, in each of which the weight of the tradable sector, ρ, is roughly around 0.4.

(14)

$$cpi = 0.4 \cdot \{0.533\,ulc - 0.329\,openness + 0.003\,t\} + 0.6 \cdot \{0.566\,ulc + 0.026\,stock\ price + 0.003\,t\}$$

$$= 0.553ulc - 0.132openness + 0.016stock\ price + 0.003t$$

As comparisons, it would be interesting to ask how the obtained long-run models, (12) and (13), based on the sectoral estimation approach, would differ from the one obtained through the traditional non-sectoral (aggregate) modeling approach. To verify this, I estimate two cointegrating VAR regressions with variables (*cpi, ulc, openness,* and *un*) and (*cpi, ulc, openness,* and *stock price*). The results of the augmented Johansen cointegration tests are reported in Table I.9.[21]

The first cointegration test with variables (*cpi*, *ulc*, *openness*, and *un*) holds well (model 1 in Table I.9). Trace statistics indicate the existence of at most one cointegrating relationship. All β coefficients satisfy sign conditions and statistically significant. However, in comparison with model (13), the size of a β coefficient on *ulc* is slightly smaller, while β coefficients on *openness* and *un* are greater. Parameter restriction tests show that α coefficients in *ulc* and *un* equations can be zero—likelihood ratio tests cannot reject the null hypothesis of $\alpha_{21} = 0$ and $\alpha_{31} = 0$ with a probability of 0.327 (restricted cointegration model 1 in Table I.9). The obtained model is:

(15) $$cpi = \underset{(0.044^{**})}{0.543}\,ulc - \underset{(0.052^{**})}{0.224}\,openness - \underset{(0.006^{**})}{0.018}\,un + \underset{(0.000^{**})}{0.004}\,t$$

with the corresponding α vector,

[21] The results of VAR lag selection are reported in Appendix Table I.3.

$$\begin{matrix} cpi \\ ulc \\ openness \\ un \end{matrix} \begin{matrix} \cdots \\ \cdots \\ \cdots \\ \cdots \end{matrix} \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \\ \alpha_{41} \end{pmatrix} = \begin{pmatrix} -0.149^{**} \\ 0.000 \\ -0.412^{**} \\ 0.000 \end{pmatrix}.$$

The robustness of this model is verified using a recursive regression methodology (Figure I.11). The coefficient on *ulc* shifts slightly in 1998-2000 but is stable in the other periods.

For the second cointegration test with endogenous variables (*cpi*, *ulc*, *openness*, and stock price), trace statistics indicate the existence of at most one cointegrating relationship (model 2 in Table I.9). However, a β coefficient on *stock price* is far from statistically significant, leading to the conclusion that this specification does not hold. Thus, it is not possible to replicate a model with the same specification as equation (14) based on the traditional non-sectoral modeling approach.

Thus, there is an interesting finding. Traditional and sectoral approaches do not lead to a same result. For the case of Japanese inflation data in this paper, the sectoral approach suggests a possibility of two cointegrating VAR models with four endogenous variables (*cpi*, *ulc*, *openness*, and *un*) and (*cpi*, *ulc*, *openness*, and *stock price*), while the traditional approach suggests only the former model with (*cpi*, *ulc*, *openness*, and *un*). Which approach suits better for Japanese data would be an empirical question, which I address in the next sub section.

E. Phillips Curve Markup Type Inflation Functions Including Short-run Dynamics

Combining the above long-run relationships with short-run dynamics, inflation functions are derived below. As a recap, the long-run equations to be embedded in inflation functions are as following.

Summary of Long Run Equations

Long-run equations 1 (*LR1*) Obtained through the sectoral approach with *openness* and *un* as markup factors (equation 13)	$cpi - (0.639ulc - 0.132openness - 0.007un + 0.003t)$
Long-run equation 2 (*LR2*): Obtained through the sectoral approach with *openness* and *stock price* as markup factors (equation 14)	$cpi - (0.553ulc - 0.132openness + 0.016stock\ price + 0.003t)$
Long-run equation 3 (*LR3*): Obtained through traditional (aggregation) approach (equation 15)	$cpi - (0.543\,ulc - 0.224\,openness - 0.018\,un + 0.004\,t)$

Figures I.12 and 13 plot the above three long-run markup relationships. Although there are some differences, the overall movement of these measures is broadly similar: markup was low and fell to a bottom of the sample periods around 1973, followed by a steady increase until the early 1980s. Markup has since fluctuated cyclically, but with much smaller variations than in the 1970s.

A general-to-specific approach is employed to select the most parsimonious model. The model selection process starts with a very general model, including a long-run relationship, LR_{t-1}, and short-run dynamics, $dcpi_{t-i}$, $dulc_{t+1-i}$, and $dipe_{t+1-i}$, where *i* takes 1 to 4.[22] In addition to these variables, both types of models include gap_{t+1-i}, oil shock dummies (*oil74q1*, 1974Q1=1; *oil74q2*, 1974Q2=1; and *oil80q1*, 1980Q1=1), high growth period dummies (*highgrowth*, 1971Q1-2003Q3=1), zero interest rate period dummies (*zeroint*, 1999Q1-2006Q1=1), consumption tax dummies (*tax89*, 1989Q2=1 and *tax97*, 1997Q2=1)— which account for the introduction of 3 percent consumption tax in April 1989 and its subsequent increase to 5 percent in April 1997— and constant and trend terms. This initial model can be considered as very general and congruent to the data evidence. Then, statistically insignificant variables are dropped with diagnostic tests to verify the validity of reduction to ensure the congruency of the final model. The final models I have obtained are:[23]

Inflation model 1 with LR1:

$$dcpi_t = \underset{[4.920]}{0.393} - \underset{[-4.908]}{0.077}\, LR1_{t-1} + \underset{[6.122]}{0.405}\, dcpi_{t-1} + \underset{[2.886]}{0.142}\, dcpi_{t-2} - \underset{[-3.991]}{0.186}\, dcpi_{t-4} + \underset{[5.050]}{0.119}\, dulc_t + \underset{[2.552]}{0.062}\, dulc_{t-1}$$
$$+ \underset{[3.867]}{0.083}\, dulc_{t-2} + \underset{[5.183]}{0.001}\, gap_{t-1} - \underset{[-4.766]}{0.001}\, gap_{t-3} + \underset{[12.293]}{0.041}\, oil74q1 - \underset{[5.318]}{0.025}\, oil74q2 + \underset{[3.739]}{0.011}\, tax89 + \underset{[3.795]}{0.011}\, tax97$$
$$- \underset{[-4.035]}{0.010}\, highgrowth$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.11; AR 1-4 test (p value),:0.10; ARCH 1-4 test (p value), 0.10; and Hetero test (p value), 0.04.

[22] "*d*" denotes the first difference operator.

[23] I used *PcGets* for calculations.

Inflation model 2 with LR2:

$$dcpi_t = \underset{[4.088]}{0.318} - \underset{[-4.0.76]}{0.064}\, LR2_{t-1} + \underset{[6.296]}{0.426}\, dcpi_{t-1} + \underset{[2.652]}{0.137}\, dcpi_{t-2} - \underset{[-3.446]}{0.162}\, dcpi_{t-4} + \underset{[4.896]}{0.119}\, dulc_t + \underset{[2.407]}{0.061}\, dulc_{t-1}$$
$$+ \underset{[3.682]}{0.081}\, dulc_{t-2} + \underset{[4.661]}{0.001}\, gap_{t-1} - \underset{[-3.942]}{0.001}\, gap_{t-3} + \underset{[11.070]}{0.040}\, oil74q1 - \underset{[-5.866]}{0.029}\, oil74q2 + \underset{[3.523]}{0.010}\, tax89 + \underset{[3.825]}{0.011}\, tax97$$
$$- \underset{[-3.520]}{0.0101}\, highgrowth$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.18; AR 1-4 test (p value),:0.03; ARCH 1-4 test (p value), 0.92; and Hetero test (p value), 0.03.

Inflation model 3 with LR3:

$$dcpi_t = \underset{[3.669]}{0.266} - \underset{[-3.656]}{0.050}\, LR2_{t-1} + \underset{[6.328]}{0.433}\, dcpi_{t-1} + \underset{[3.015]}{0.155}\, dcpi_{t-2} - \underset{[-3.370]}{0.164}\, dcpi_{t-4} + \underset{[4.795]}{0.117}\, dulc_t + \underset{[2.559]}{0.065}\, dulc_{t-1}$$
$$+ \underset{[3.821]}{0.085}\, dulc_{t-2} + \underset{[4.569]}{0.001}\, gap_{t-1} - \underset{[-4.541]}{0.001}\, gap_{t-3} + \underset{[11.545]}{0.042}\, oil74q1 - \underset{[-5.721]}{0.028}\, oil74q2 + \underset{[3.508]}{0.010}\, tax89 + \underset{[3.894]}{0.011}\, tax97$$
$$- \underset{[-3.116]}{0.009}\, highgrowth$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.08; AR 1-4 test (p value), 0.84; ARCH 1-4 test (p value), 0.75; and Hetero test (p value), 0.02.

where [ ] indicated *t* statistics.

The main findings are that all measures of long-run markup relationships are highly statistically significant and have negative coefficients, consistent with the predicted error correction mechanism. Besides, incidentally, all these models include the same set of short-run and dummy variables. Lags of *dcpi* and *dulc* are highly significant variables. While the energy price (*dipe*)— the level of which was previously found insignificant in the long-run model— is again not significant, two oil shock dummies are positive (in the summation of these two) and significant. This suggests that the first oil shock significantly affected

inflation in Japan, but for the other periods, imported energy prices did not matter. Furthermore, all these models are reasonably parsimonious, as the total number of parameters has been reduced from 28 in the initial model to 15 in the final model. The results of the diagnostic tests are overall satisfactory, although all models fail in heteroschedasticity tests (model 2 also fails in AR1-4 test).

Note that while the first and fourth lags of *gap* are significant, combining them together, their contribution to inflation appears to be none. This would suggest that rather than the level of *gap*, its change, $dgap_t = gap_t - gap_{t-1}$, would be important. To test hypothesis, I re-estimate inflation model 1 including gap_{t-1} and $dgap_{t-1}$, and obtain the following:

Revised inflation model 1 with LR1:

$$dcpi_t = \underset{[4.591]}{0.382} - \underset{[-4.580]}{0.075}\, LR1_{t-1} + \underset{[7.287]}{0.491}\, dcpi_{t-1} + \underset{[0.970]}{0.048}\, dcpi_{t-2} - \underset{[-3.270]}{0.156}\, dcpi_{t-4} + \underset{[5.006]}{0.122}\, dulc_t + \underset{[2.414]}{0.064}\, dulc_{t-1}$$
$$+ \underset{[2.705]}{0.056}\, dulc_{t-2} + \underset{[0.968]}{0.0001}\, gap_{t-1} + \underset{[3.560]}{0.001}\, dgap_{t-1} + \underset{[11.659]}{0.041}\, oil74q1 - \underset{[-5.647]}{0.028}\, oil74q2 + \underset{[3.749]}{0.011}\, tax89 + \underset{[4.077]}{0.012}\, tax97$$
$$- \underset{[-3.496]}{0.009}\, highgrowth$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.11; AR 1-4 test (p value),:0.01; ARCH 1-4 test (p value), 0.55; and Hetero test (p value), 0.00.

Note that coefficient on gap_{t-1} becomes insignificant, while that on $dgap_{t-1}$ is highly significant, supporting the hypothesis that a change in output gap would be more important than its level. Similar conclusions hold for the other two inflation models with LR2 and LR3, as well (not reported). Accordingly, dropping gap_{t+1-i} and adding $dgap_{t+1-i}$ where *i*

takes 1 to 4, I estimate these three models all over again following the general-to-specific approach, and obtain the following.

Inflation model 1 with LR1 (final):

$$dcpi_t = \underset{[4.981]}{0.378} - \underset{[-4.969]}{0.074}\, LR1_{t-1} + \underset{[5.995]}{0.417}\, dcpi_{t-1} + \underset{[2.356]}{0.130}\, dcpi_{t-2} - \underset{[-3.933]}{0.185}\, dcpi_{t-4} + \underset{[5.286]}{0.121}\, dulc_t + \underset{[2.829]}{0.069}\, dulc_{t-1}$$
$$+ \underset{[3.924]}{0.082}\, dulc_{t-2} + \underset{[2.829]}{0.001}\, dgap_{t-1} + \underset{[3.206]}{0.001}\, dgap_{t-2} + \underset{[12.525]}{0.041}\, oil74q1 - \underset{[5.266]}{0.025}\, oil74q2 + \underset{[3.679]}{0.011}\, tax89 + \underset{[3.899]}{0.011}\, tax97$$
$$- \underset{[-4.001]}{0.010}\, highgrowth$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.10; AR 1-4 test (p value),:0.08; ARCH 1-4 test (p value), 0.76; and Hetero test (p value), 0.01.

Inflation model 2 with LR2 (final):

$$dcpi_t = \underset{[4.109]}{0.288} - \underset{[-4.0.95]}{0.058}\, LR2_{t-1} + \underset{[6.152]}{0.438}\, dcpi_{t-1} + \underset{[2.224]}{0.127}\, dcpi_{t-2} - \underset{[-3.360]}{0.159}\, dcpi_{t-4} + \underset{[5.162]}{0.122}\, dulc_t + \underset{[2.813]}{0.070}\, dulc_{t-1}$$
$$+ \underset{[3.856]}{0.084}\, dulc_{t-2} + \underset{[3.610]}{0.001}\, dgap_{t-1} + \underset{[2.810]}{0.001}\, dgap_{t-2} + \underset{[11.710]}{0.041}\, oil74q1 - \underset{[-5.740]}{0.029}\, oil74q2 + \underset{[3.463]}{0.010}\, tax89 + \underset{[3.956]}{0.011}\, tax97$$
$$- \underset{[-3.427]}{0.010}\, highgrowth$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.14; AR 1-4 test (p value),:0.04; ARCH 1-4 test (p value), 0.90; and Hetero test (p value), 0.002.

Inflation model 3 with LR3:

$$dcpi_t = \underset{[3.756]}{0.267} - \underset{[-3.742]}{0.051}\, LR3_{t-1} + \underset{[6.202]}{0.445}\, dcpi_{t-1} + \underset{[2.462]}{0.141}\, dcpi_{t-2} - \underset{[-3.262]}{0.159}\, dcpi_{t-4} + \underset{[4.834]}{0.114}\, dulc_t + \underset{[2.620]}{0.066}\, dulc_{t-1}$$
$$+ \underset{[3.718]}{0.082}\, dulc_{t-2} + \underset{[3.719]}{0.001}\, dgap_{t-1} + \underset{[2.844]}{0.001}\, dgap_{t-2} + \underset{[11.589]}{0.041}\, oil74q1 - \underset{[-5.595]}{0.028}\, oil74q2 + \underset{[3.452]}{0.010}\, tax89 + \underset{[3.894]}{0.011}\, tax97$$
$$- \underset{[-3.143]}{0.009}\, highgrowth$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.10; AR 1-4 test (p value), 005; ARCH 1-4 test (p value), 0.82; and Hetero test (p value), 0.002.

where [ ] indicated *t* statistics. The final models include the first and second lags of *dgdp*. These models remain parsimonious with all variables highly significant. Residual tests are also generally satisfactory, except for heteroschedasticity tests.

So, which Phillips-curve type model best explains the data generating process of inflation in Japan? To address this question, I employ an encompassing methodology to compare the above three models. The general concept of encompassing principle is that a null model H_0 is considered to encompass an alternative model H_1, if it can explain the results of the latter model. Among several encompassing test statistics, I use the *J test* proposed by Davidson and McKinnon (1993). Their idea is that if one model is the correct one, then the fitted values from the other model should not have explanatory power in estimating that model. For example, we are interested in comparing the following two competing linear models:

$H_0 : y = X\beta + \varepsilon_0$, and $H_1 : y = Z\gamma + \varepsilon_1$,

where y is a $n \times 1$ matrix, X and Z are $n \times k$ matrices, and β and γ are $k \times 1$ matrixes.

To test a null model, H_0, against an alternative model, H_1, first, estimate model H_1 to obtain fitted values $\hat{y} = Z\hat{\gamma}$. Then, estimate the following model.

$$y = X\beta + \hat{y}\delta + \varepsilon .$$

$$= X\beta + (Z\hat{\gamma})\delta + \varepsilon$$

If δ (scalar) is significant, model H_0 can be rejected, while if δ (scalar) is insignificant, model H_0 can be accepted. Similarly, to test model H_1 against H_0, first, estimate model H_0 to get fitted values $\hat{y} = X\hat{\beta}$. Then, estimate the following model.

$$y = Z\gamma + (X\hat{\beta})\theta + \varepsilon_1 .$$

If θ (scalar) is significant, model H_1 can be rejected, or vice versa. Finally, if both H_0 and H_1 are either significant or insignificant, the model comparison is inconclusive. The data do not provide enough information to discriminate between the two models.

Table I.10 reports the results of the encompassing tests. The fitted value of model 2 is insignificant in model 1 with the *p* value of 0.30, while the fitted value of model 1 is significant in model 2 at the 1 percent significance level. Similarly, the fitted value of model 3 is also insignificant in model 1 with the *p* value of 0.07, while the fitted value of model 1 is highly significant in model 3.[24] This suggests that model 1 encompasses both models 2 and 3. Between models 2 and 3, the former appears to weakly encompass the latter because the fitted value of model 3 is far from significant in model 2 with the *p* value of 0.79, while the fitted value of model 2 is insignificant in model 3 at a much lower *p* value of 0.11. The overall conclusion is that model 1 encompasses model 2, which in turn

[24] Appendix Figure 1 shows the results of recursive estimates (starting 1985Q4) of the encompassing tests between models 1 and 3. The figure shows *t* statistics associated with the fitted values. It suggests that the fitted values of model 1 in model 3 (null hypothesis) has been constantly highly significant (Appendix Figure 1 bottom). On the contrary, the level of significance of the fitted values of model 3 in model 1 (null hypothesis) has been steadily falling and has crossed the 5 percent significance level since 2001 (Appendix

encompasses model 3. Thus, the two models derived from the sectoral modeling approach perform better than the traditional approach.

II. Monetary type inflation modeling

As an alternative approach to Phillips curve markup type modeling, a monetary approach is considered. Since Friedman advocated that "inflation is the monetary phenomena," money has been considered to be an important determinant to long run prices. In empirical literature, one of the common approaches to incorporating money into an inflation model is estimating a real money gap, which is defined as the difference between actual real stock of money and its long run level: see Hallman, Porter, and Small (1991), Juselius (1992), Altimari (2001), Gerlach and Svensson (2000), and Sekine (2001).

A. Modeling Strategy

Define the real money gap as following:

(16) $real\ money\ gap = rm_t - rm_t^*$,

where rm_t is the level of current real money stock and rm_t^* is its long-run level. Following the earlier studies, a long-run real money stock, rm_t^*, is defined as:

(17) $rm_t^* = k_y^* y_t - k_{oc}^* \cdot oc_t + k_a^* \cdot a_t$,

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Figure I.1 Top). Accordingly, I conclude that model 1 encompasses model 3 stronger than vice versa.

where y_t is output (real GDP), oc_t is the opportunity costs of holding money, and a_t is a variable capturing a wealth effect. The coefficients, k_y^*, k_i^*, and k_a^*, are the long run elasticities of income, opportunity costs, and wealth effect, respectively. Equation (17) can be interpreted as a modified Keynesian type money demand function, including a wealth effect. These three long-run elasticities are obtained by estimating a long run real money demand function.[25]

B. Key Data

[25] Following Hallman, Porter, and Small (1991) and Gerlach and Svensson (2000), *real money gap* can also be interpreted in the context of a p-star model. From the quantity theory of money, define the long-run steady-state real money stock, rm_t^{SS}, as:

$$rm_t^{SS} \equiv y_t^* - v_t^* ,$$

where y_t^* is the level of potential output and v_t^* is the long run velocity. rm_t^{SS} is the level of real money stock consistent with the level of potential output and long-run velocity. This equation means that the long-run steady-state real money stock, rm_t^{SS}, is the level of real money stock consistent with the level of potential output and long-run velocity. rm_t^{SS} can also be interpreted as the level of long-run real money stock rm_t^* evaluated at a steady state. Then, define p_t^* as the price level that would make the value of the current money stock equal to the real long-run money stock with output at its potential, y_t^*, and velocity at its long-run level, v_t^*.

$$p_t^* \equiv m_t - rm_t^{SS} .$$

From these two equations,

$$\begin{aligned} p_t^* - p_t &\equiv (m_t - rm_t^{SS}) - p_t \\ &= rm_t - rm_t^{SS} \end{aligned}$$

Accordingly, the real money gap is identical to the negative of the price gap, $p_t - p_t^*$. If the current price is below (above) p-star, $p_t^* - p_t > 0$ ($p_t^* - p_t < 0$), there is excess money (shortage of money) in the economy and price is expected to increase (decrease).

All data are quarterly (seasonally adjusted) and available for 1970Q1-2007Q1. The new data that require explanation include money stock, opportunity costs of holding money, and proxy for wealth, while data for GDP and prices are those used in the previous section of Phillip curve markup type modeling. For money stock, M2CD—broad monetary including CD—is chosen because this has been seen as a policy reference monetary indicator of the Bank of Japan for a long period.[26] The opportunity cost of holding money is defined as the difference between 10-year government bond yields minus money own rates (weighted average of interest rates of assets in M2CD).[27] As proxy for a wealth variable, Nikkei 225 stock exchange index (*real stock price*) and the estate stock price index (*real estate stock price*), both deflated by CPI, are used. Table I.11 summarizes the descriptive statistics of these new variables.

Figures I.14-16 plot the logarithm (except for *oc*) of these key variables. Both real M2CD (*rm2cd*) and real GDP (*rgdp*) follow similar logarithmic trends, although they seem to have been slowly diverging since the early 1980's. The opportunity cost, *oc* (in percent), and wealth variables, *real stock price* and *real estate stock price* are clearly much more volatile, and it is not immediately clear whether there exists a long-run relationship between *rm2cd* and these variables.

[26] Source: Research and Statistics Department, The Bank of Japan, http://www.boj.or.jp/en/theme/research/stat/money/ms/index.htm.
[27] Source, CEIC.

The results of ADF unit root tests are reported in Table I.12. Four variables, *rm2cd*, *oc*, *real stock price*, and *real estate stock price* are clearly non-stationary I(1), as the null hypothesis of the existence of unit root cannot be rejected at the 5 percent significance level. On the contrary, the test results are ambiguous for *rgdp*: it is stationary, if a trend is excluded, while it is non stationary, if a trend is included (the inclusion of the trend is not statistically supported). However, an estimated coefficient on the lagged value, γ*-1*, is quite close to zero, and thus, as in the previous section, I employ an alternative KPSS stationarity tests. The test results show that the null hypothesis, i.e., *rgdp* is stationary, can be rejected at the 1 percent significance level, regardless with or without the inclusion of a trend. Accordingly, I conclude that *rgdp* is an I(1) variable.

C. Estimation

I estimate cointegrating VAR and obtained results as reported in Table I.13.[28] In the most simplistic form (model 1 in Table I.13), there exists a long-run cointegrating relationship among *rm2cd*, *rgdp*, and *oc*. The β coefficients on *rgdp* and *oc* show a negative and positive sign, respectively—consistent with the theoretical predictions—, and both are highly significant. All the speed of adjustment coefficients are significant, and signs of these coefficients ($\alpha_{11} < 0, \alpha_{21} > 0$, and $\alpha_{31} < 0$) support an error correction mechanism. The size of the β coefficient on *rgdp* is 1.23, larger than unity, which implies that the accumulation of financial assets has been faster than GDP growth in Japan.

[28] The results of VAR lag selection are reported in Appendix Table I.4.

Adding *real estate stock price*, a four variable cointegrating VAR indicates the existence of one long-run relationship (model 2). β coefficient on *rgdp* remains virtually unchanged from model 1 and is highly significant. β coefficient on *oc* is smaller but remains significant. A new variable, *real estate stock price*, is also highly significant, suggesting a positive relation between *real estate stock price* and *real money demand*. On the speed of adjustment coefficients, all α coefficients, except on the *real estate stock price*, equation are significant. Accordingly, I test a parameter restriction, $\alpha_{41} = 0$ and find that the likelihood ratio test cannot reject the null hypothesis with the probability of 0.26 (model 2 with restricted cointegration).

Replacing *real estate stock price* with *real stock price*, four variable cointegrating VAR regression suggests the existence of one cointegration relationship. However, β coefficient on *real stock price* is insignificant. Accordingly, this model is dropped from the subsequent analysis.

Based on the above results, I choose model 2 with the parameter restriction, $\alpha_{41} = 0$.[29]

(18) $rm2cd = \underset{(0.034)}{1.277}\, rgdp - \underset{(0.008)}{0.034}\, oc + \underset{(0.013)}{0.044}\, a$,

with the corresponding α vector,

[29] The size of the β coefficient on *rgdp* remains greater than unity at 1.277. To verify the robustness of this result, I tested the joint null hypothesis of β coefficient on *rgdp =0* and $\alpha_{41} = 0$. The test was rejected at the probability value of 0.00.

$$\begin{matrix} rm2cd \\ rgdp \\ oc \\ real\,estate\,stock\,price \end{matrix} \begin{pmatrix} \dots \\ \dots \\ \dots \\ \dots \end{pmatrix} \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \\ \alpha_{31} \end{pmatrix} = \begin{pmatrix} -0.086^{**} \\ 0.045^{*} \\ -4.630^{**} \\ 0.000 \end{pmatrix}.$$ [30]

Thus, long run income, opportunity costs, and wealth effect elasticities are $k_y = 1.277$, $k_{oc} = 0.034$, and $k_a = 0.044$, respectively. Recursive regressions confirm that all the β coefficients are highly stable (Figures I.17).

D. Monetary Type Inflation Functions Including Short-run Dynamics

From equation 18, the *real money gap* can be expressed as:

Long-run equation 4 (*LR4*): $rm2cd - (1.277\,rgdp - 0.034\,oc + 0.044\,a)$

Figures I.18 plots *real money gap*. *Real money gap* shows large volatilities, with three peaks in the early 1970s, late 1990s, and early 2000s most prominent. The first peak corresponds to an excessive monetary expansion just before the first oil shock, which led to a very high inflation in Japanese monetary history after the post World War II periods. The second peak corresponds to BOJ's policy easing in the fear of economic recessions following sharp exchange rate appreciations after the Plaza and Louvre Accords—and later on, Japan experienced so called boom and bust of bubbles. The last peak represents BOJ's efforts during the periods of zero interest rate and quantitative monetary easing policy.

[30] This cointegration holds for a shorter sample period, 1981Q1-2007Q1 with similar

A very general monetary type inflation function is estimated including the lag of *real money gap*, $LR4_{t-1}$, and short-run dynamics, $dcpi_{t-i}$ and $dipe_{t-i}$ where *i* takes 1 to 4. In addition, a model includes $dgap_{t+1-i}$,[31] and exogenous variables are the same as the previous Phillips curve markup type model (*oil74q1*, *oil74q2*, *oil80*, *tax89*, *tax97*, *high growth period*, *zero interest period*, and constant and trend terms). This general model is reduced by eliminating irrelevant variables following the general-to-specific principle. The final model I have obtained is:

Inflation Model 4 with LR4:

$$dcpi_t = \underset{[3.558]}{0.191} + \underset{[3.553]}{0.030}\, LR4_{t-1} + \underset{[8.408]}{0.562}\, dcpi_{t-1} + \underset{[4.087]}{0.179}\, dcpi_{t-2} + \underset{[5.713]}{0.257}\, dcpi_{t-3} - \underset{[-3.007]}{0.132}\, dcpi_{t-4}$$
$$+ \underset{[2.643]}{0.001}\, dgap_{t-1} + \underset{[13.812]}{0.045}\, oil74q1 - \underset{[-5.195]}{0.025}\, oil74q2 + \underset{[4.586]}{0.014}\, tax89 + \underset{[5.247]}{0.016}\, tax97$$
$$+ \underset{[2.554]}{0.003}\, highgrowth - \underset{[-2.755]}{0.002}\, zeroint$$

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.01; AR 1-4 test (p value), 0.00; ARCH 1-4 test (p value), 0.29; and Hetero test (p value), 0.001.

The main conclusions derived from the above estimation results are that *real money gap*, *LR4*, is highly significant, and with its positive sign, a larger real money gap implies a greater pressure on inflation. Lags of *dcpi* are also highly significant. However, imported energy prices (*dipe*) are again not significant, although a sum of the oil shock dummies are positive. This suggests that while a major oil shock in the early 1970s raised inflation, in

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parameter estimates.

[31] Initially, the level of *gap* was also included, together with its difference. However, similar to the Phillips curve type inflation models discussed earlier, I found that the level of *gap* was not significant, and thus was dropped.

the other periods, energy prices did not matter a lot in Japanese inflation. Similar to the case of Phillips curve type inflation model 1, a change in output gap, $dgap_{t-1}$, affects inflation. Finally, unfortunately, diagnostic tests are not fully satisfactory, with very low statistics for *normality*, *AR1-4*, and *heteroschedasticity* tests.

V. Model Comparisons—Phillips Curve Markup Type Versus Monetary Type Inflation Functions

Empirical evidence suggests that both the Phillips curve markup type inflation function and the monetary type inflation function explain the data generation process of Japanese inflation reasonably well. As a recap, the inflation modes selected thus far are as following.

Summary of Inflation Functions

Phillips curve markup type inflation model (model 1), with LR1 obtained through the sectoral approach with *openness* and *un* as markup factors	$dcpi_t = \underset{[4.981]}{0.378} - \underset{[-4.969]}{0.074}\, LR1_{t-1} + \underset{[5.995]}{0.417}\, dcpi_{t-1} + \underset{[2.356]}{0.130}\, dcpi_{t-2} - \underset{[-3.933]}{0.185}\, dcpi_{t-4} + \underset{[5.286]}{0.121}\, dulc_t + \underset{[2.829]}{0.069}\, dulc_{t-1} + \underset{[3.924]}{0.082}\, dulc_{t-2} + \underset{[2.829]}{0.001}\, dgap_{t-1} + \underset{[3.206]}{0.001}\, dgap_{t-2} + \underset{[12.525]}{0.041}\, oil74q1 - \underset{[5.266]}{0.025}\, oil74q2 + \underset{[3.679]}{0.011}\, tax89 + \underset{[3.899]}{0.011}\, tax97 - \underset{[-4.001]}{0.010}\, highgrowth$
Monetary type inflation model (model 4), with the real money gap (LR4)	$dcpi_t = \underset{[3.558]}{0.191} + \underset{[3.553]}{0.030}\, LR4_{t-1} + \underset{[8.408]}{0.562}\, dcpi_{t-1} + \underset{[4.087]}{0.179}\, dcpi_{t-2} + \underset{[5.713]}{0.257}\, dcpi_{t-3} - \underset{[-3.007]}{0.132}\, dcpi_{t-4} + \underset{[2.643]}{0.001}\, gap_{t-1} + \underset{[13.812]}{0.045}\, oil74q1 - \underset{[-5.195]}{0.025}\, oil74q2 + \underset{[4.586]}{0.014}\, tax89 + \underset{[5.247]}{0.016}\, tax97 + \underset{[2.554]}{0.003}\, highgrowth - \underset{[-2.755]}{0.002}\, zeroint$

But which one explains better? To answer this question, I compare models 1 and 4 using encompassing tests (Table I.14). The table shows that the null hypothesis of model 1 against model 4 (i.e., model 1 encompasses model 4) can be rejected at the 1 percent significance level, while the null hypothesis of Inflation Model 4 against Inflation Model 1 can also be rejected at the 1 percent significant level. Thus, encompassing test results are inconclusive.[32]

This result may suggest that it would be worth attempting to model an inflation function including both markup and real money gap. To this end, I first estimate a general model including both LR_{t-1} (*markup*) and $LR4_{t-1}$ (real money gap) as long-run relationships, and $dcpi_{t-i}$, $dulc_{t+1-i}$, $dipe_{t+1-i}$, and $dgap_{t+1-i}$ as short-run factors, where *i* takes 1 to 4. The exogenous variables are the same as before. As in the previous case, the goal is to get a parsimonious and data congruent model following a general-to-specific approach. A final model that I have obtained is:

(19)

$$\begin{aligned} dcpi_t = {} & \underset{[5.195]}{0.440} - \underset{[-4.992]}{0.074}\, LR1_{t-1} + \underset{[1.612]}{0.010}\, LR4_{t-1} + \underset{[6.100]}{0.422}\, dcpi_{t-1} + \underset{[2.362]}{0.129}\, dcpi_{t-2} - \underset{[-3.796]}{0.178}\, dcpi_{t-4} \\ & + \underset{[5.286]}{0.120}\, dulc_t + \underset{[2.801]}{0.068}\, dulc_{t-1} + \underset{[3.991]}{0.084}\, dulc_{t-2} + \underset{[4.040]}{0.001}\, dgap_{t-1} + \underset{[3.046]}{0.001}\, gap_{t-2} + \underset{[12.307]}{0.041}\, oil74q1 - \underset{[-5.428]}{0.026}\, oil74q2 \\ & + \underset{[3.708]}{0.011}\, tax89 + \underset{[4.148]}{0.012}\, tax97 - \underset{[-4.186]}{0.010}\, highgrowth \end{aligned}$$

[32] However, the results of recursive estimates (starting 1985Q4) support stronger evidence of model 1 encompassing model 4 (Appendix Figure I.2). The figure shows *t* statistics associated with the fitted values. The fitted values of model 4 in model 1 (null hypothesis) has become significant just several years ago (Appendix Figure I.2 bottom), while the level of significance of the fitted values of model 1 in model 4 (null hypothesis) has been steadily high (Appendix Figure I.2 top).

Sample period: 1971Q2-2007Q1; S.E.=0.003; Normality test (p value), 0.10; AR 1-4 test (p value), 0.12; ARCH 1-4 test (p value), 0.85; and Hetero test (p value), 0.02.

The first long run relationship, *LR1* (markup), is highly significant. The second long-run relationship, *LR2* (real money gap), has a theory consistent positive sigh, but is significant only at the 10 percent level. To further verify this finding, I estimate the above model recursively. The results suggest that coefficients on *LR1* (markup) have been constantly highly significant, while the significant level of those on *LR2* (real money gap) have remained relatively low (Figure I.19). In sum, I conclude that there is some evidence that the former Phillips curve markup type model better explain the Japanese inflation process.

Figures I.20 and 21 plot actual and fitted values, residuals, and the results of recursive estimates of the Inflation Model 1 and 4. Both models track actual outcomes fairly closely. The results of recursive estimates—1 step residuals, 1 step Chow tests, and break point Chow tests—broadly suggest the stability of the models, although both models show a large spike in 1 step residual and 1 step Chow tests in the third quarter of 1989. This irregularity likely reflects the impact of the introduction of the consumption tax in the second quarter of 1989, which has not been fully captured by the use of tax dummy, *tax 89*.

VI. Policy Implications

Recently, there are signs that CPI inflation has begun to increase. It appears that deflationary pressures peaked out around 2002. CPI inflation has rising back to and around

zero percent, albeit at a very gradual pace. In the final quarter of 2007, CPI inflation rose above zero percent (year-on-year) and has reached 1 percent in the past year (Figure I.2).

I am able to explain the recent inflation developments using an estimated Phillips curve-type model. Note that over the last few years, the unemployment rate has steadily declined and stock price index gone up, while the openness indicator has continued to rise (Figure I.7). The estimated long-run Phillips curve-type model (e.g., equation 13) suggests that the lower unemployment rate and the higher stock price index should have contributed to reduce *LR1*, while the higher openness should have contributed to raise *LR1*. The former factor outweighed the latter factor, resulting in a decline in markup (smaller LR1) between 2004 and 2007 as shown in Figures I.12 and 13. Recall that the estimated Phillips curve-type inflation model including short-run dynamics (model 1 with LR1) suggests that the α adjustment coefficient attached to *LR1* is negative at -0.074. Accordingly, a decrease in *LR1* should have contributed to raise inflation rate (*dcpi*). In addition, the level of GDP gap has started rising since 2002, implying positive *dgap*. This should have also contributed to raise inflation through short-run dynamics.

Unfortunately, the monetary type model provides limited insights on how to analyze the recent inflation developments. Figure I.18 shows a declining trend of real money gaps since 2002. The smaller real money gap would contribute to drive inflation rates down with the positive α adjustment coefficient attached to *LR4* (model 4 with LR4). This is a counterintuitive result.

Looking forward, the global financial crisis is expected to negatively affect the economic outlook for Japan. Since the eruption of the global financial crisis in late 2008, the stock price index in Japan has fallen significantly back to the level in 2003, a 5-year low. The current economic outlook is increasingly pessimistic, and GDP growth in Japan is projected to slowdown. Thus the unemployment rate is projected to rise and the GDP gap to fall. All in all, the estimated Phillips curvy-type inflation model suggests a strong likelihood of a re-emergence of deflationary pressures in the Japanese economy.

VII. Conclusions

My research revisits inflation functions in Japan. Major structural changes in the macro economy during the 1990s seem to have affected the long-run price determination in Japan. I focus on macroeconomic indicators which have changed significantly during the 1990s—namely, a sharp rise in openness, a persistent increase in unemployment, and a steady decline in asset prices— to examine whether these factors have affected the inflation process in Japan.

One of the main findings is that the Phillips curve markup type model explains the Japanese inflation process successfully. The rising openness has pushed downward the price level in the tradable sector, while a steady increase in unemployment and a fall in stock prices contributed to lower the price level in the non-tradable sector.

The above-finding also implies that the determinants of the inflation process in the tradable and non-tradable sectors are different. In fact, I also find that estimating a Phillips curve

inflation function separately for the tradable sector and the non-tradable sector, and then combining these functions to get the overall inflation function would be more appropriate than estimating a Phillips curve inflation function for the overall CPI.

I also find that the monetary type—real money gap—inflation model explains Japanese inflation reasonably well. The encompassing test, comparing the Phillips curve markup type inflation function and the monetary type inflation function, does not conclusively show which model dominate the other. However, if a model is estimated including both *markup* and *real money gap*, it is found that the *real money gap* is less significant than *markup*. This suggests that markup factors have affected inflation more influentially than monetary factors.

There is one major caveat to the results in my paper. Openness, unemployment, and stock prices were chosen with the prior belief that they are important in explaining a structural change. Although the data support that these variables constitute part of the long run price development, this does not mean there are no other variables affecting Japanese inflation. The identification of other possible variables will require further empirical work.

Figure I.1. CPI (Total, Tradable, and Nontradable)
(Seasonally adjusted; 2005=100)

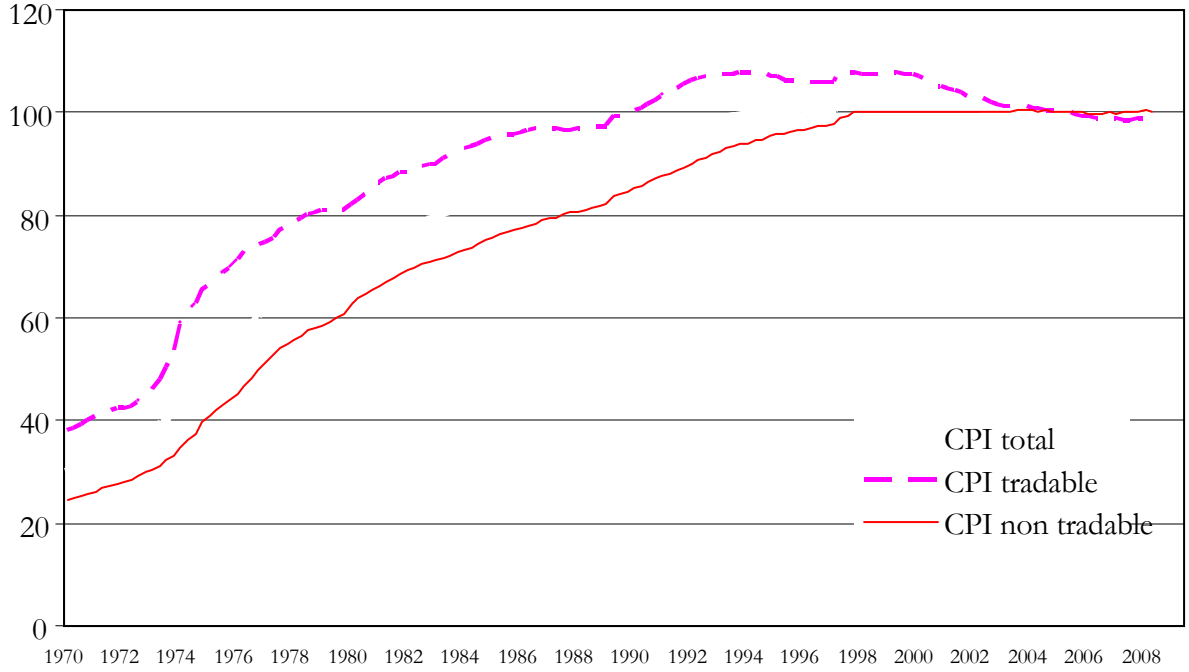


Figure I.2. CPI Inflation by Sector
(Annual percent increase)

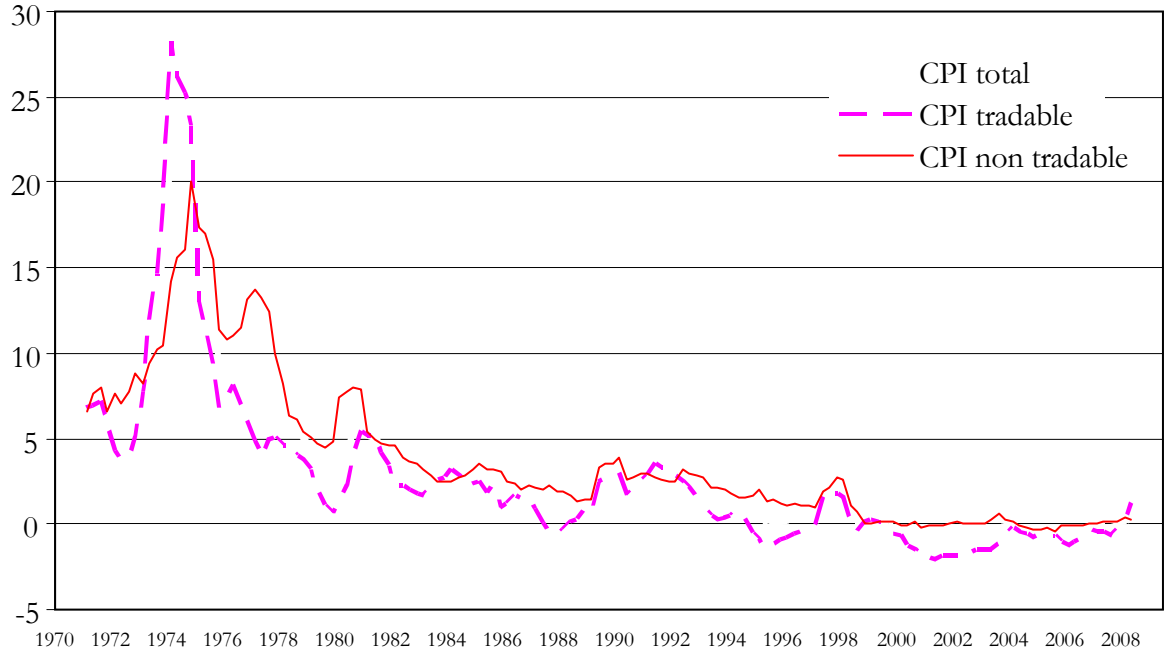


Figure I.3. Unit Labor Costs
(In logarithm)

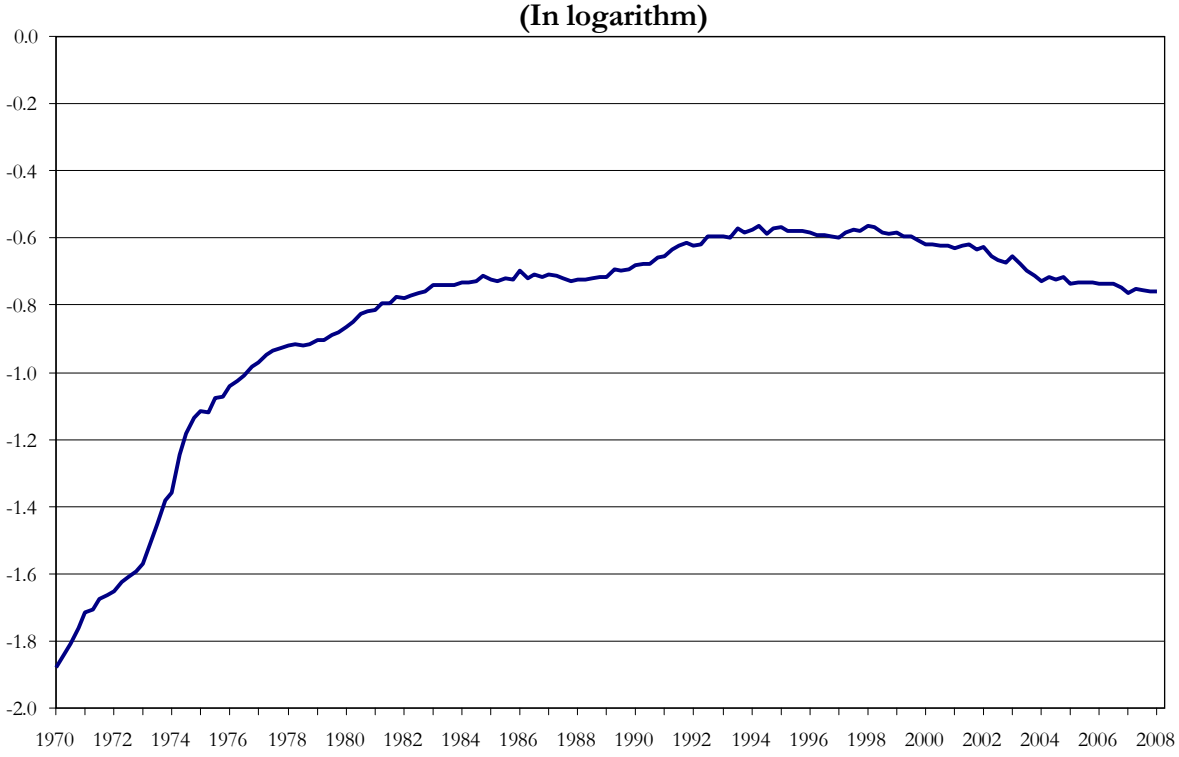


Figure I.4. Unit Labor Costs
(Annual percent change)

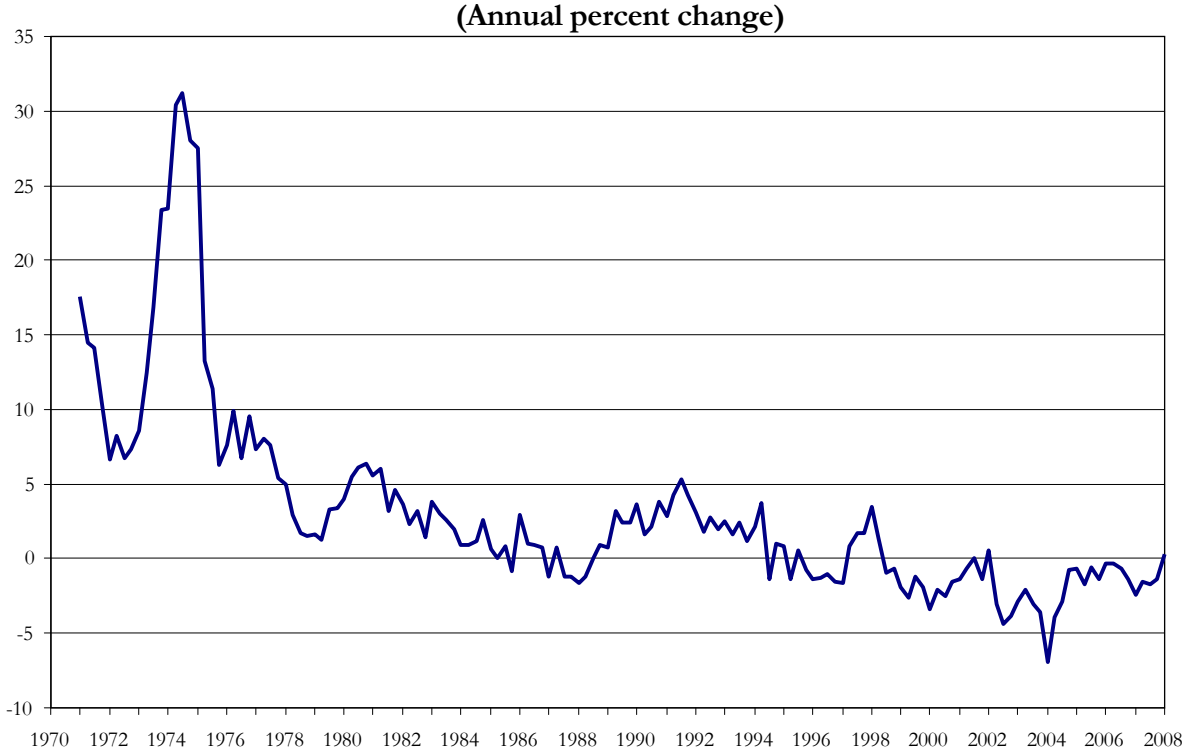


Figure I.5. Input Prices
(Seasonally adjusted; 2000=100)

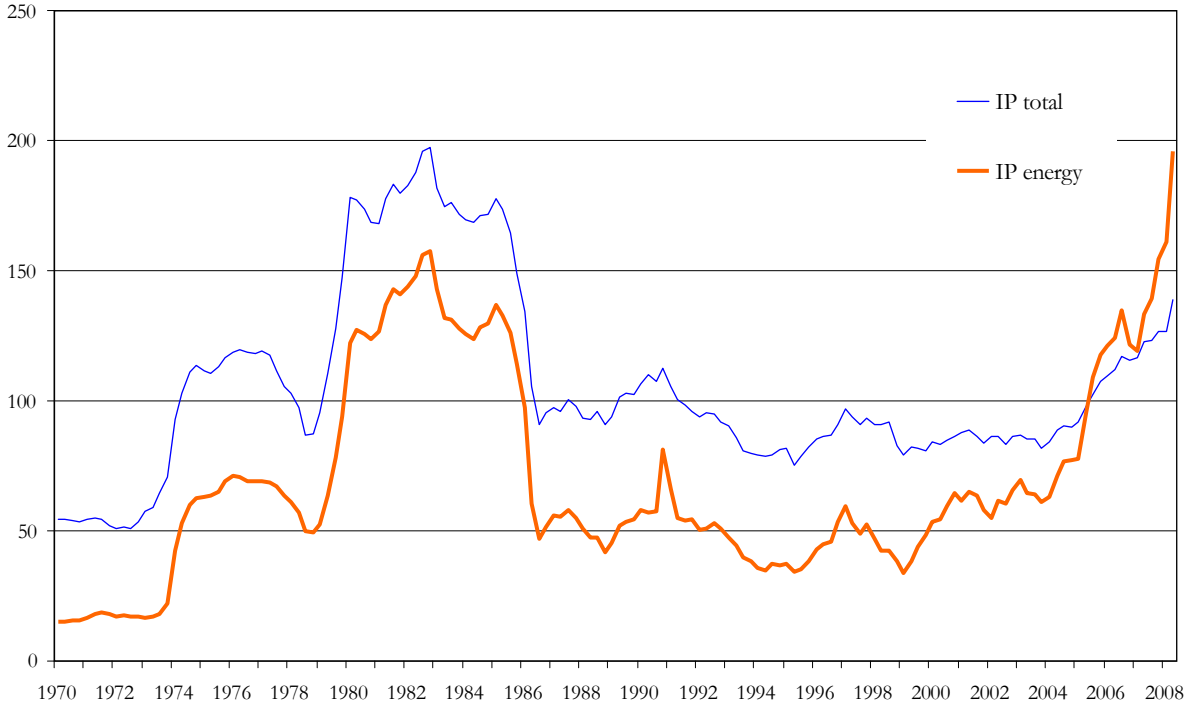
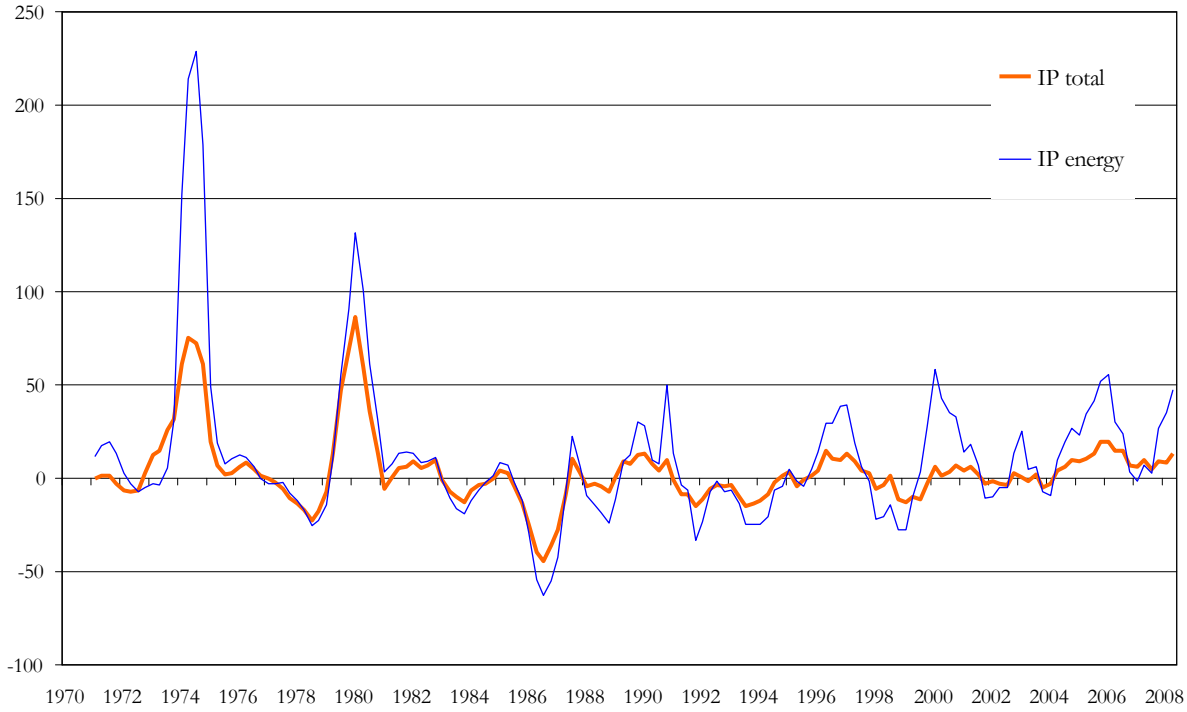
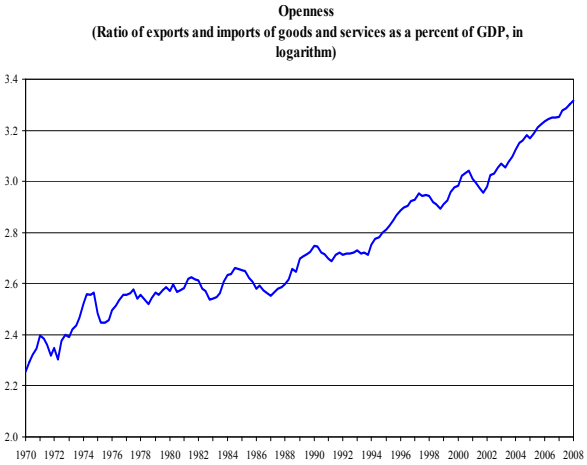
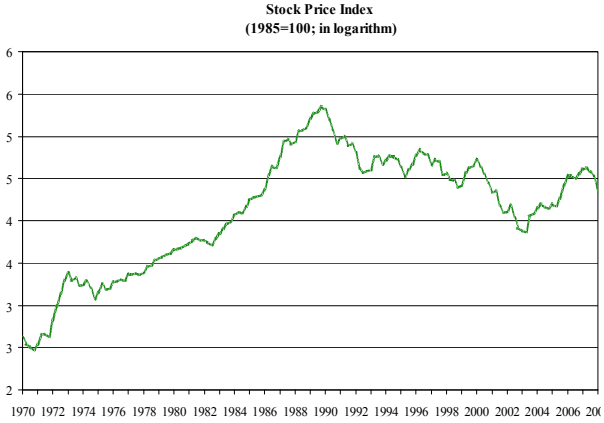


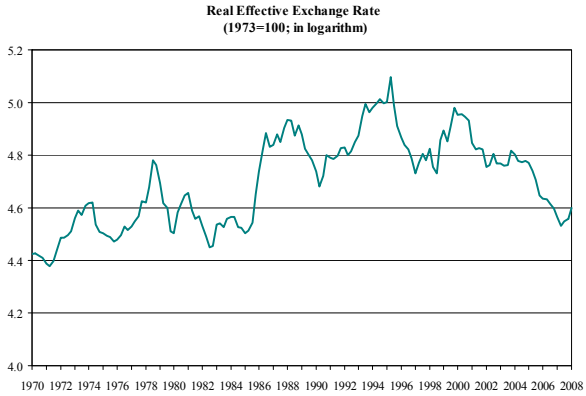
Figure I.6. Input Prices
(Annual percent increase)

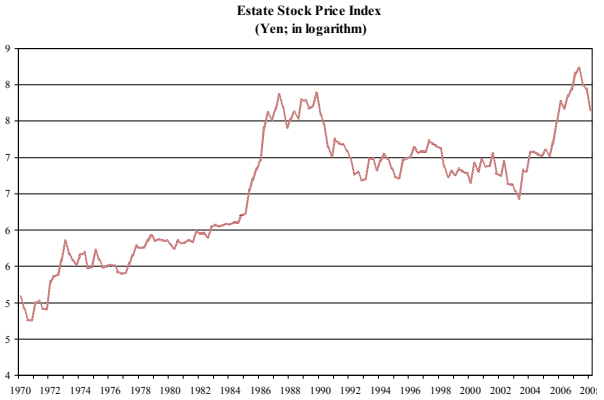


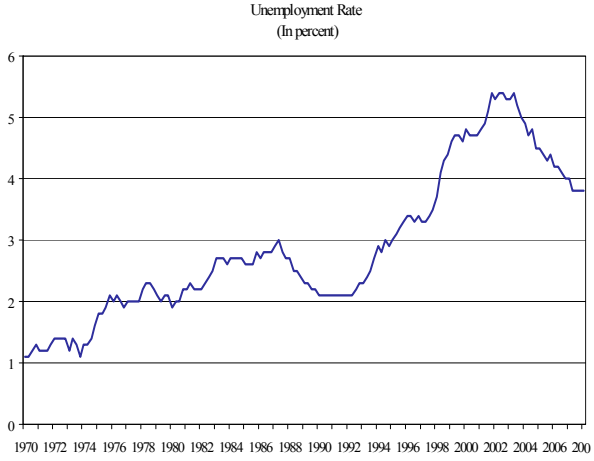
Figures I.7. Dynamic Markup Factors











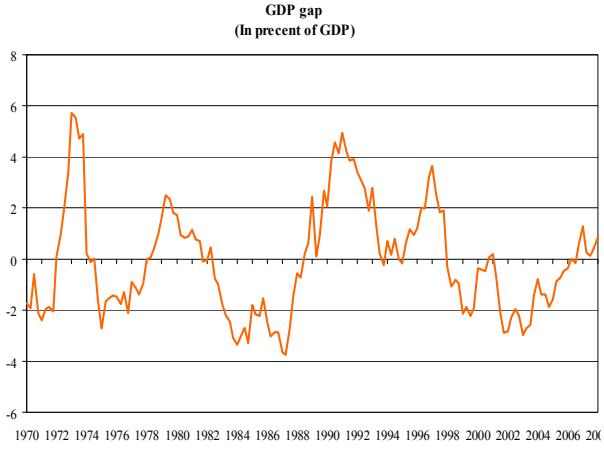
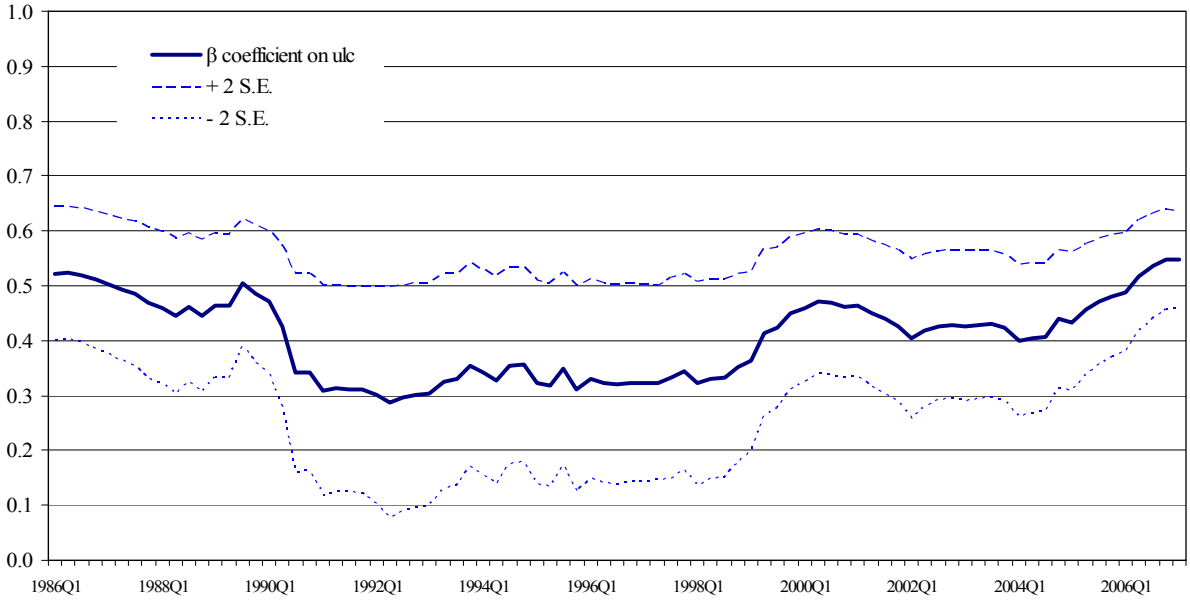


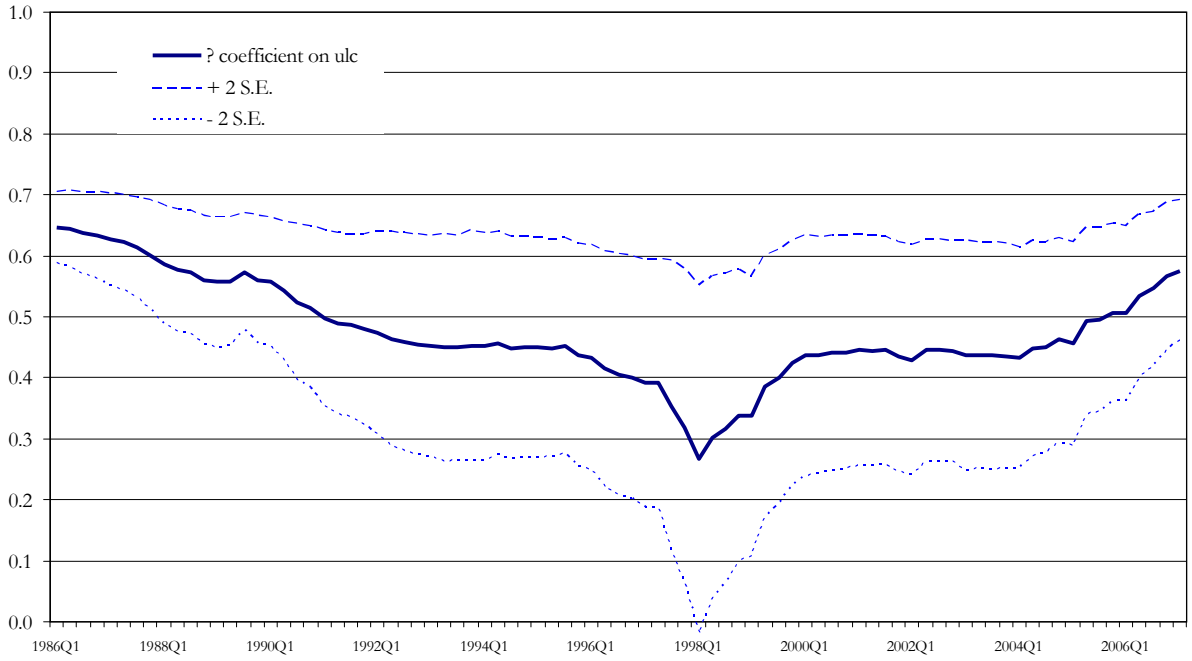
Figure I.8. Traditional Case: Phillips Curve Markup Type Inflation Model, Recursive Estimates 1/



1/ The underlying model is model 3 in Table 6.

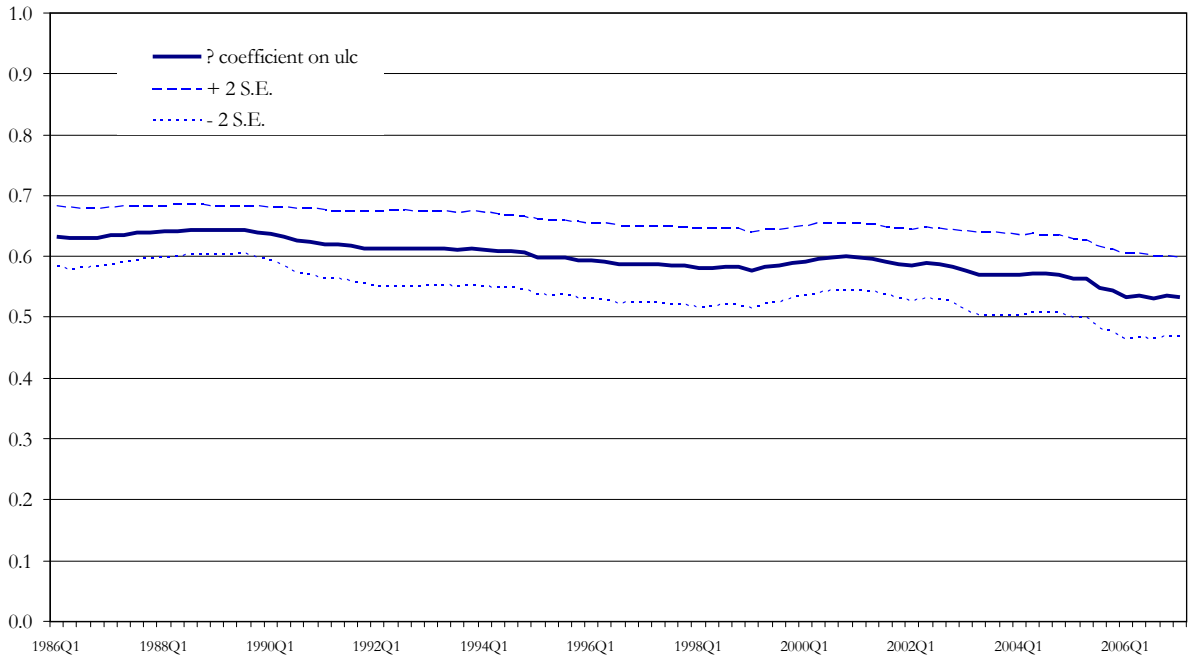
Figure I.9. Tradable Sector Phillips Curve Markup Type Inflation Model: Recursive Estimates

Figure I.9a. Tradable Sector Phillips Curve Markup Type Inflation Model with Two Endogenous Variables (cpit-ulc), Recursive Estimates 1/



1/ The underlying model is model 1 in Table 7.

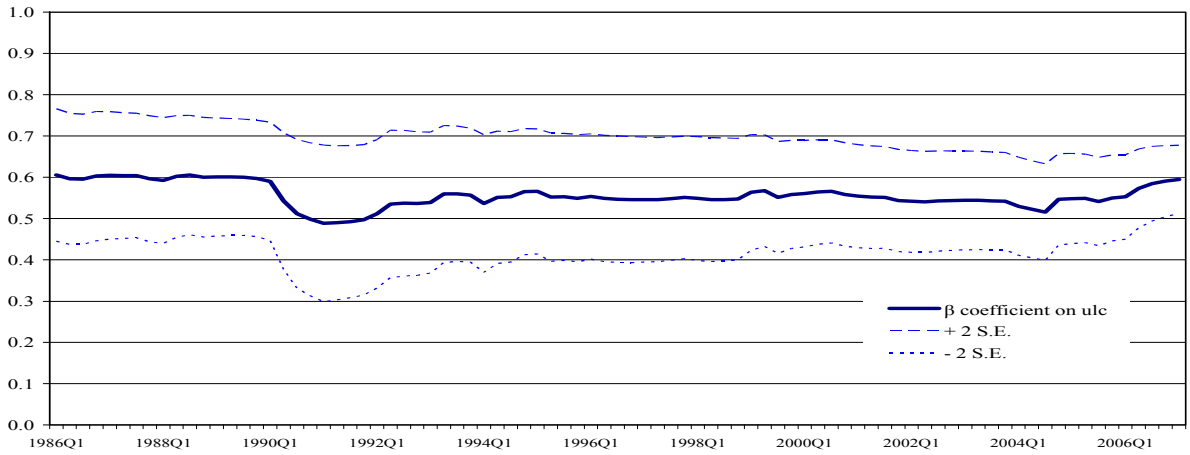
Figure I.9b. Tradable Sector Phillips Curve Markup Type Inflation Model with Three Endogenous Variables (cpit-ulc-openness): Recursive Estimates 1/



1/ The underlying model is model 4 in Table 7.

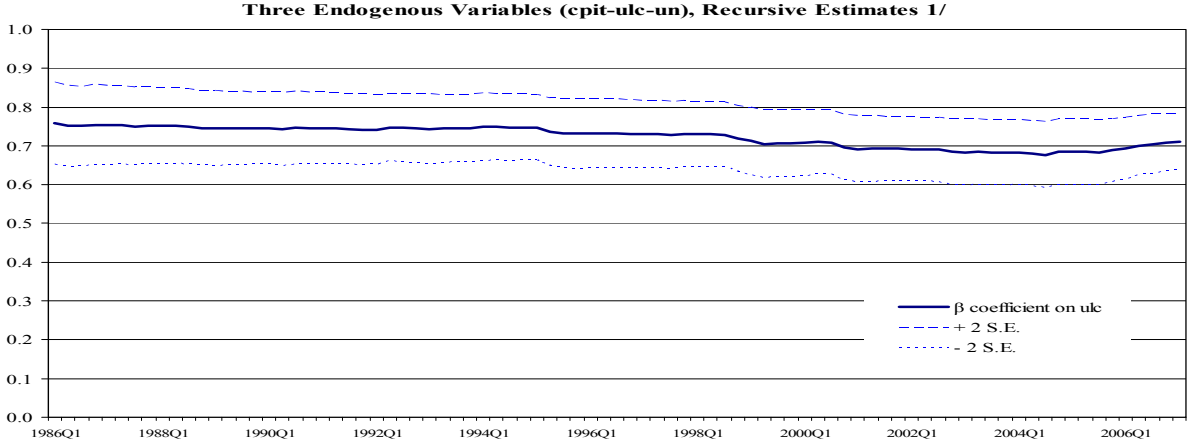
Figure I.10. Nontradable Sector Phillips Curve Markup Type Inflation Model: Recursive Estimates

Figure 10a. Nontradable Sector Phillips Curve Markup Type Inflation Model with Two endogenous Variables (cpit-ulc), Recursive Estimates 1/



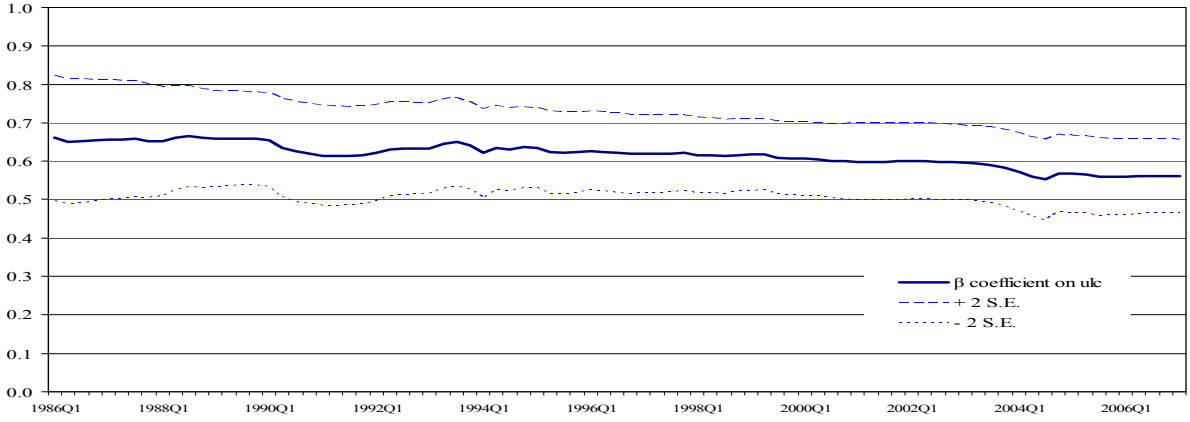
1/ The underlying model is model 1 in Table 8.

Figure 10b. Nontradable Sector Phillips Curve Markup Type Inflation Model with Three Endogenous Variables (cpit-ulc-un), Recursive Estimates 1/



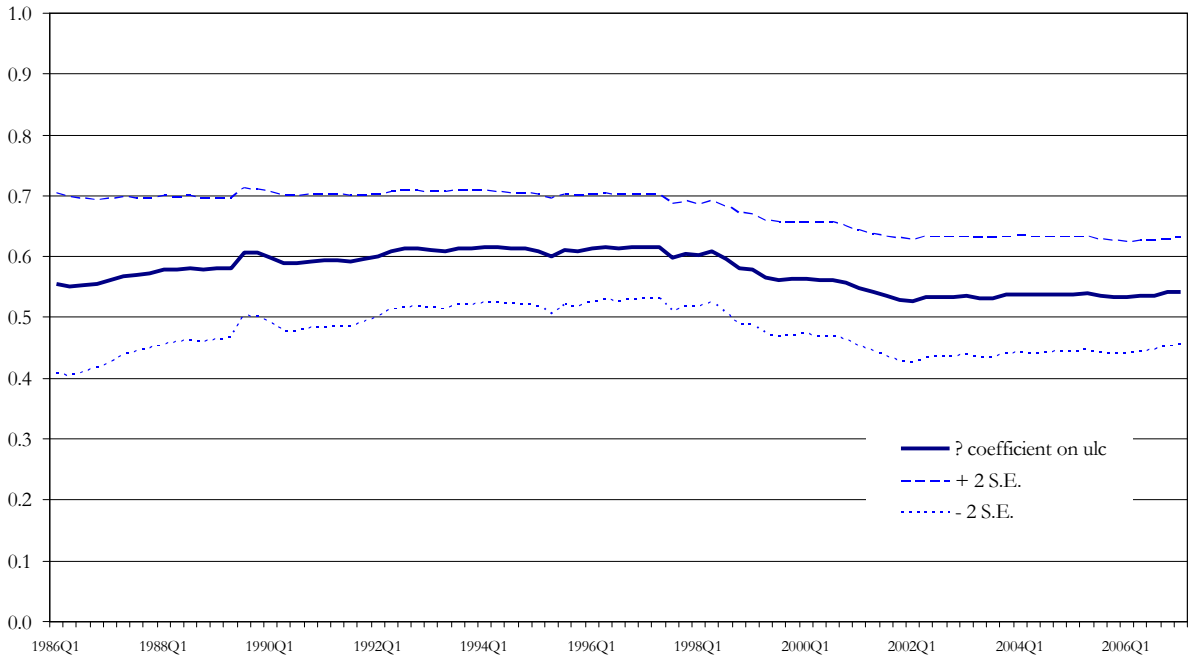
1/ The underlying model is model 4 in Table 8.

Figure 10c. Nontradable Sector Phillips Curve Markup Type Inflation Model with Three Endogenous Variables (cpit-ulc-stock price), Recursive Estimates 1/



1/ The underlying model is model 5 in Table 8.

Figure I.11. Augmented Traditional Phillips Curve Markup Type Inflation Model with Four Endogenous Variables (cpi, ulc, openness, and un): Recursive Estimates 1/



1/ The underlying model is model 1 in Table 9.

Figure I.12. Long Run Equations (Markups)
(Standarized by setting 1980Q1=1, 1970Q1-2007Q1)

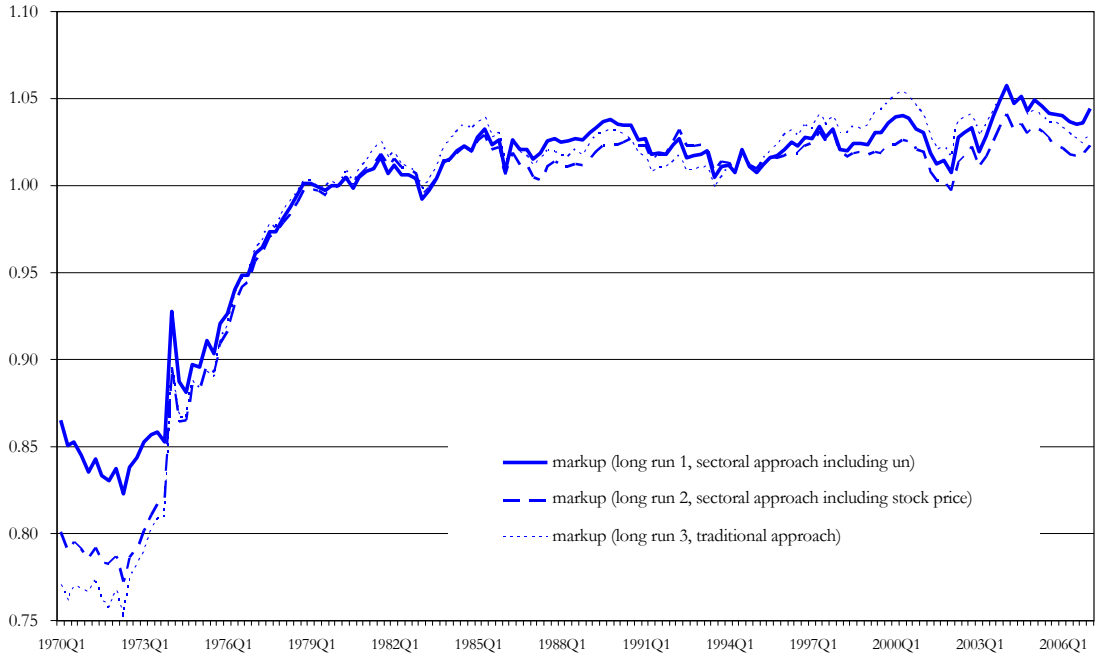


Figure I.13. Long Run Equations (Markups)
(Standarized by setting 1980Q1=1, 1980Q1-2007Q2)

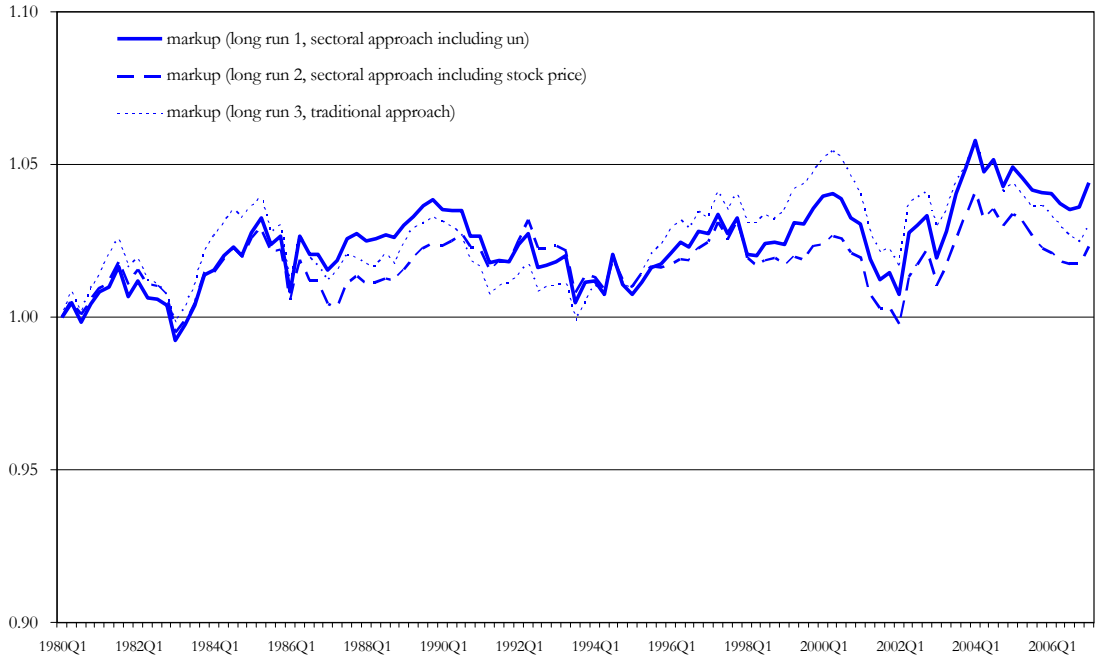


Figure I.14. Real M2CDs and real GDP
(In billions of yen in logarithm)

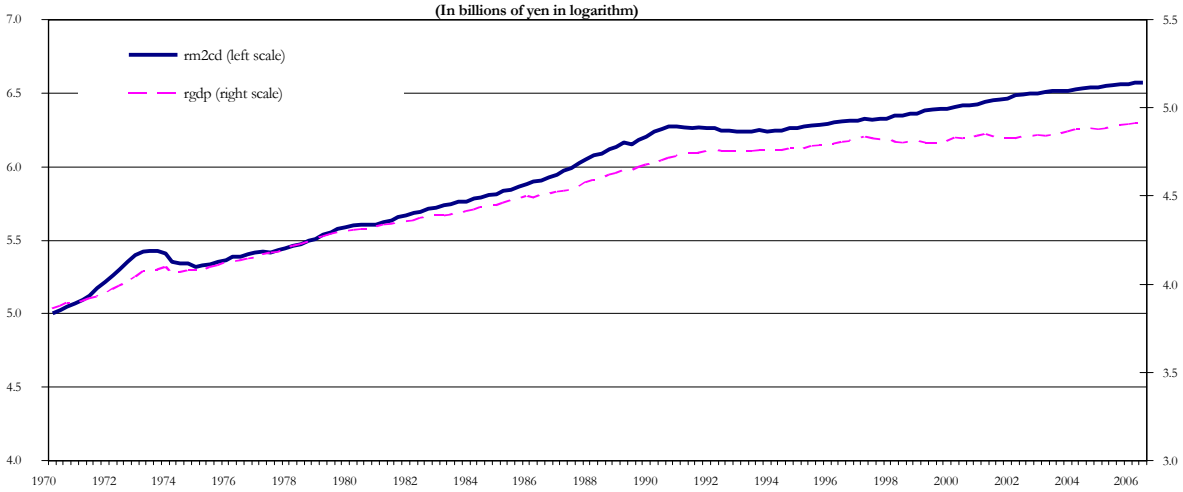


Figure I.15. Real M2CDs and Opportunity Costs
(In billions of yen in logarithm; unless otherwise indicated)

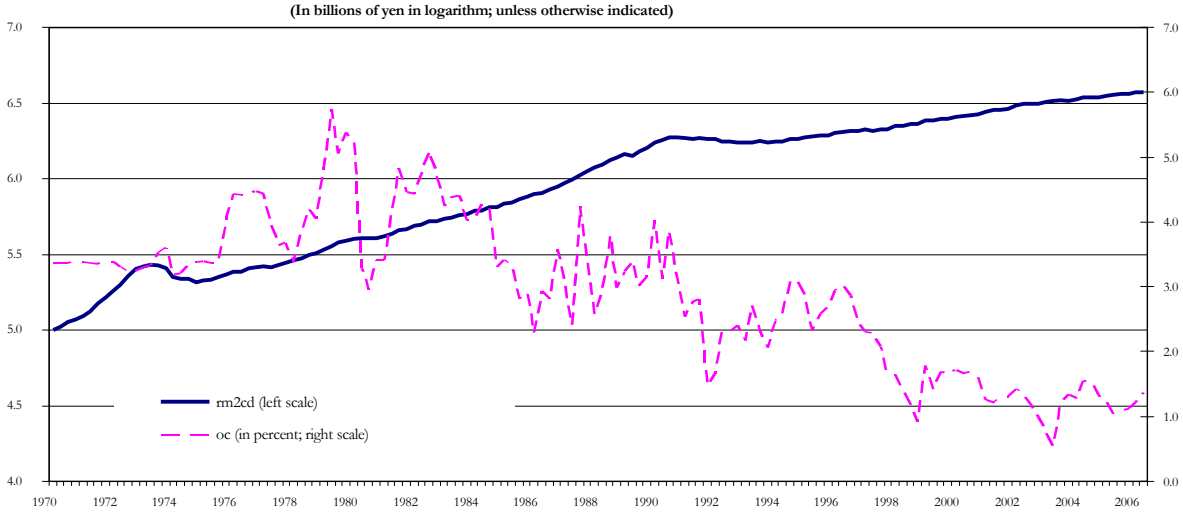
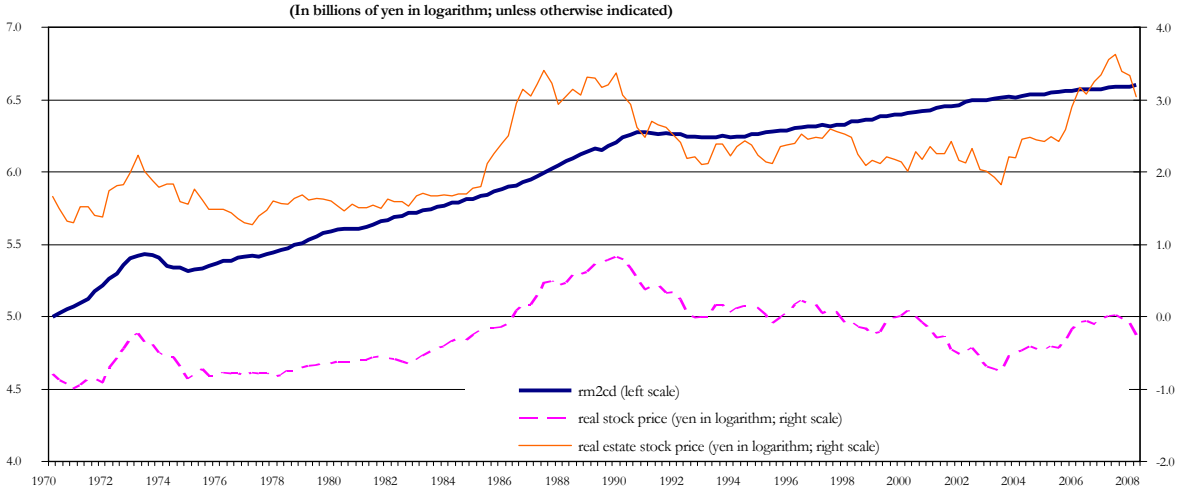
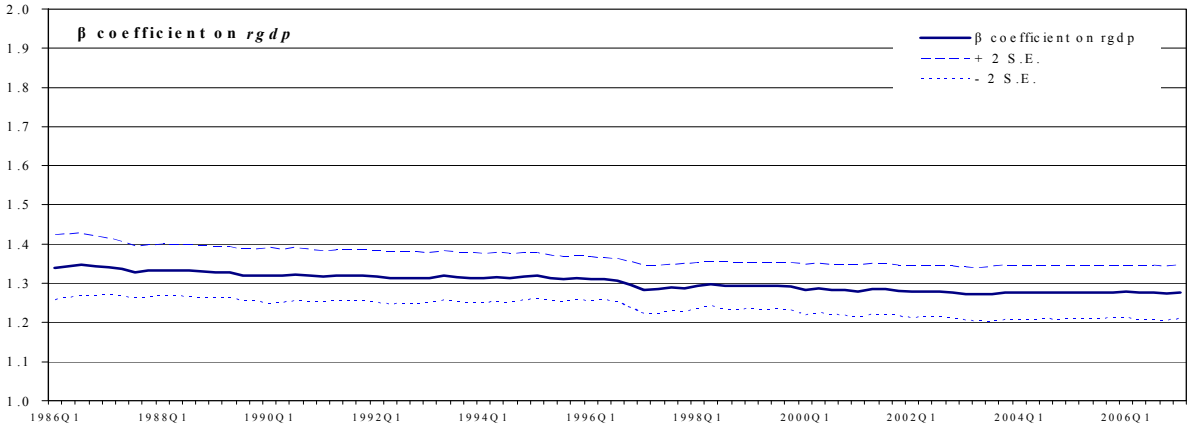


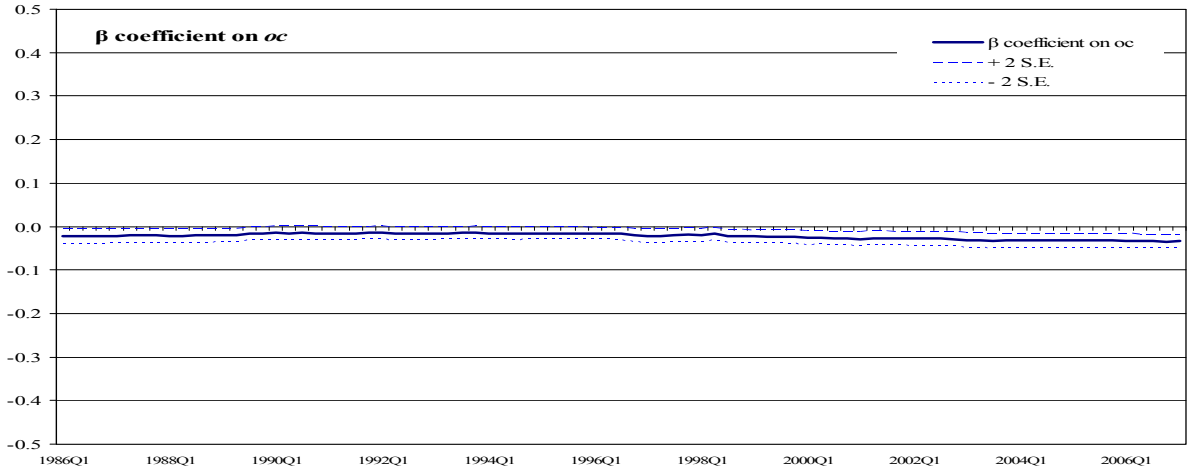
Figure I.16. Real M2CDs, Real Stock, Real Estate Stock Prices
(In billions of yen in logarithm; unless otherwise indicated)



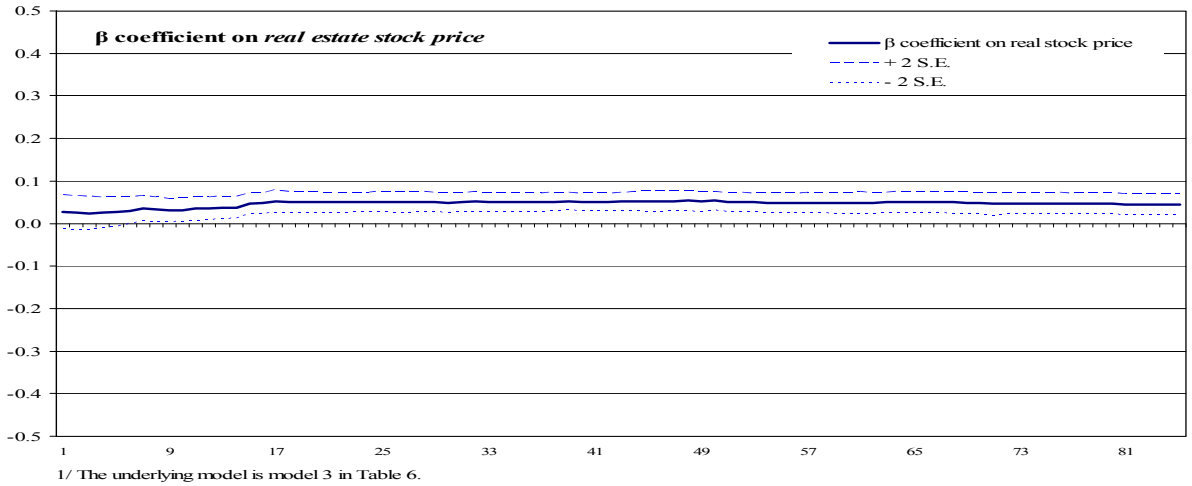
Figures I.17. Monetary Type Inflation Model (rm2cd-rgdp-oc-real estate stock price): Recursive Estimates 1/



1/ The underlying model is model 3 in Table 6.



1/ The underlying model is model 3 in Table 6.



1/ The underlying model is model 3 in Table 6.

Figure I.18. Real Money Gap
(Standarized by setting 1980Q1=1, 1970Q1-2007Q1)

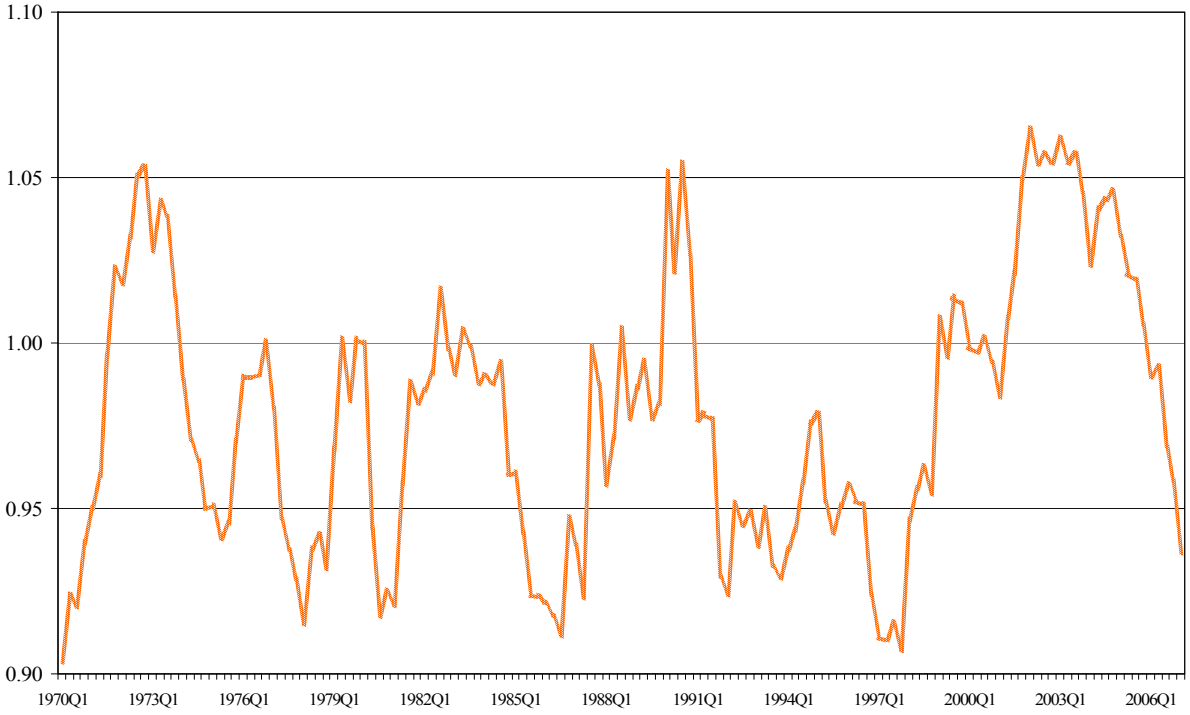


Figure I.19. Mixed Inflation Model with Long Run Relation 1 (Markup) and Long Run Relation 4 (Real Money Gap): Recursive Estimates

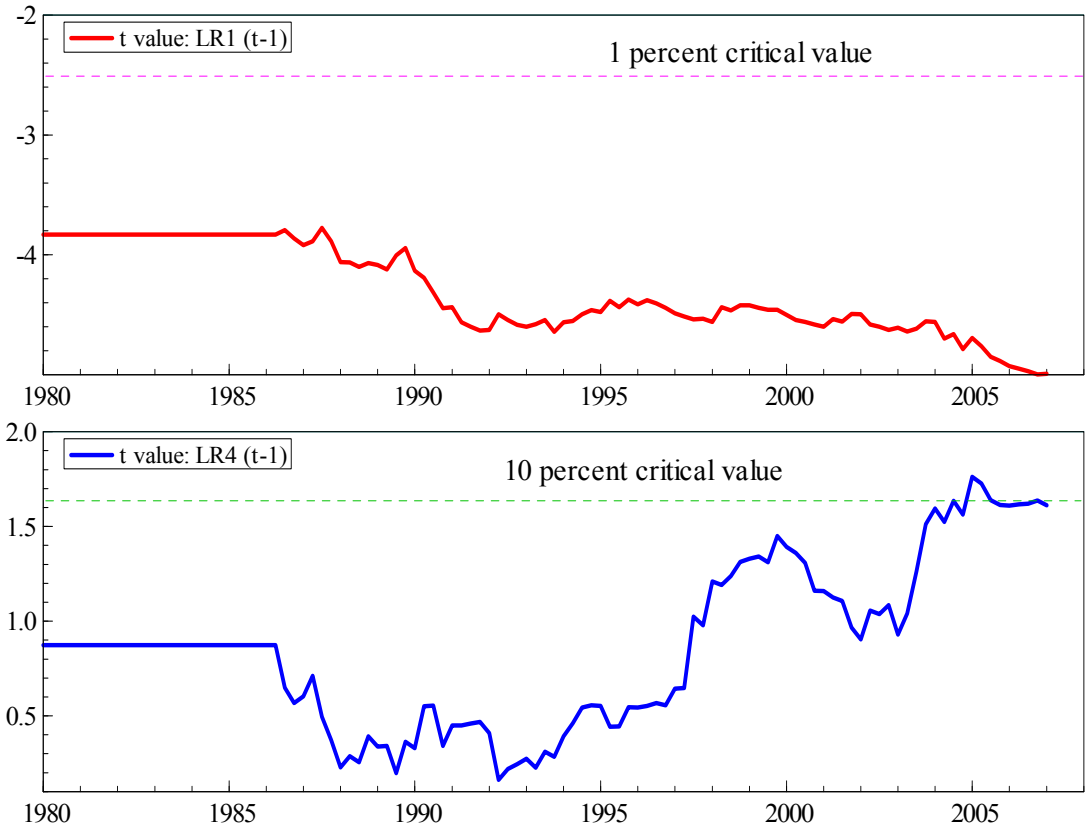


Figure I.20. Philips-curve Type Japanese Inflation Model

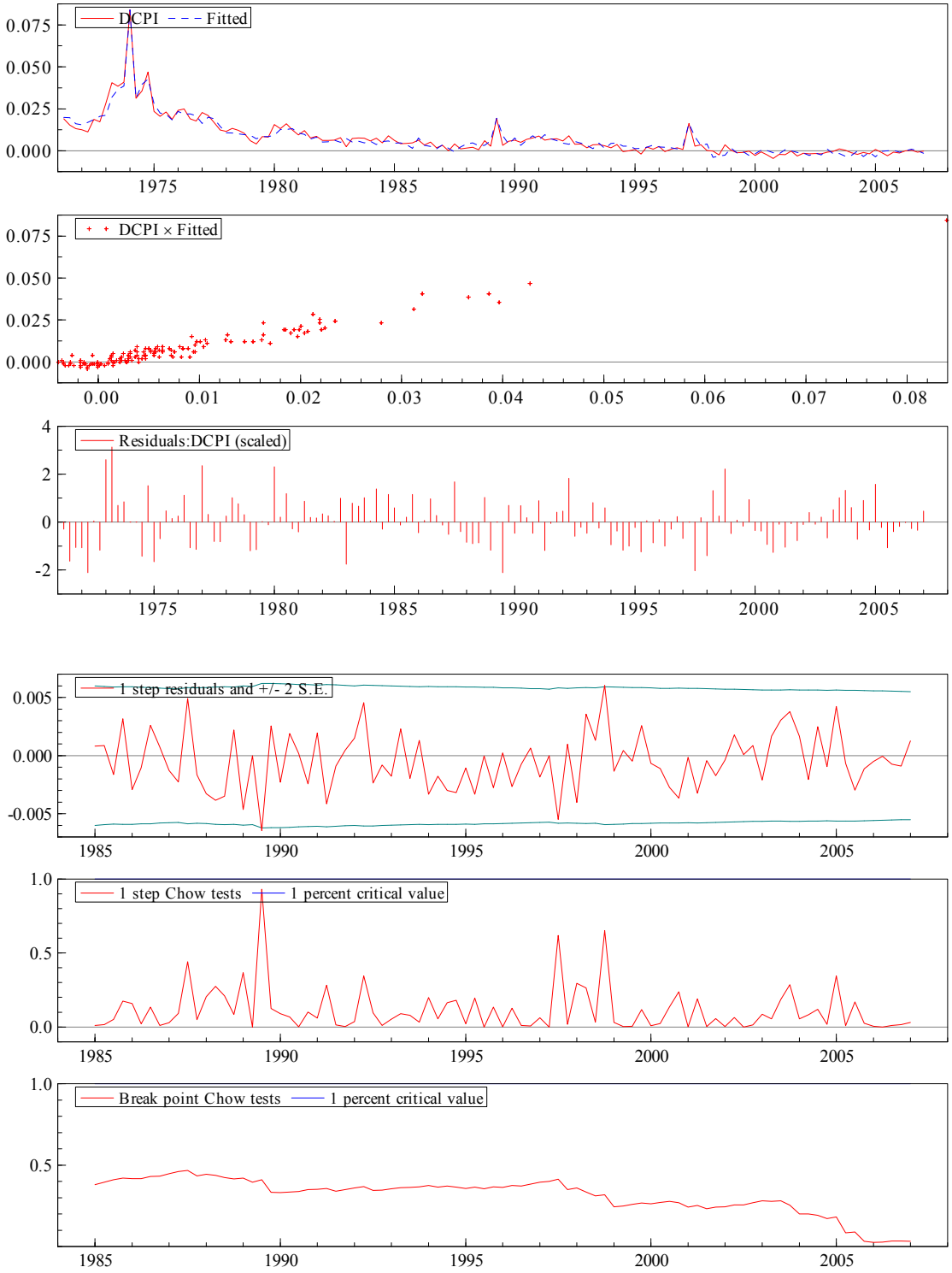


Figure I.21. Monetary Type Japanese Inflation Model

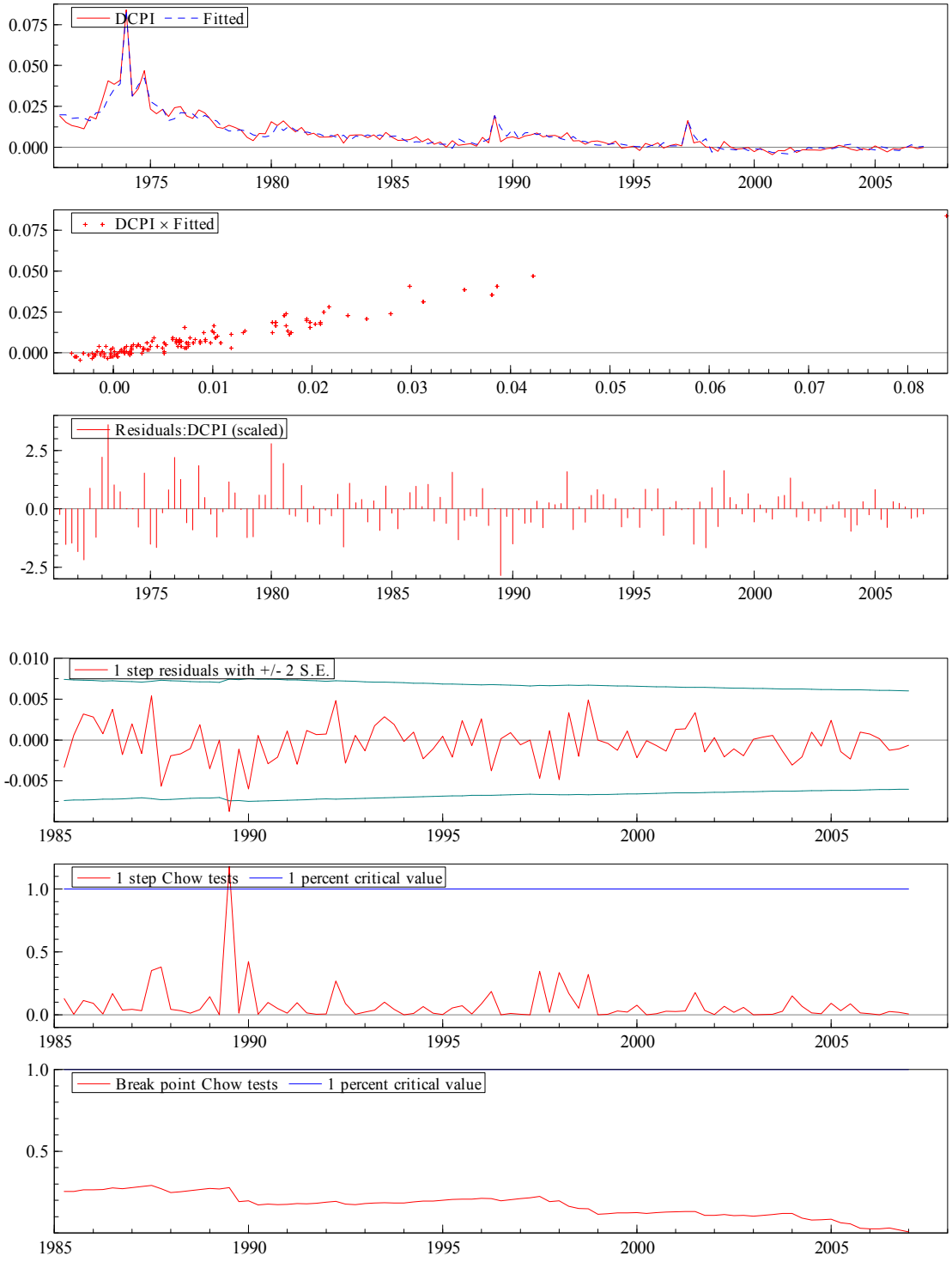


Table I.1. Phillips Curve Type Modeling: Descriptive Statistics
(In logarithm, 1970Q1 - 2007Q1)

	Mean	Standard deviation	Skewness
Level			
cpi	4.358	0.342	-1.459
cpit	4.465	0.280	-1.723
cpint	4.259	0.412	-1.252
ulc	-0.835	0.323	-1.832
ip	4.601	0.324	0.212
ipe	4.059	0.581	-0.457
openness	2.724	0.245	0.467
reer	4.705	0.171	0.005
un	2.865	1.198	0.656
stock price	4.111	0.705	-0.453
estate stock price	6.519	0.814	-0.216
gap	-0.094	2.163	0.642
First difference			
Dcpi	0.008	0.012	2.812
Dcpit	0.006	0.013	3.591
Dcpint	0.009	0.011	1.730
Dulc	0.008	0.019	1.961
Dip	0.005	0.057	0.432
Dipe	0.014	0.109	0.855
Dopenness	0.007	0.022	-0.399
Dreer	0.001	0.042	0.272
Dun	0.020	0.117	0.075
Dstock price	0.013	0.079	-0.165
Destate stock price	0.021	0.142	0.299
Dgap	0.020	0.888	-0.875

Table I.2. Phillips Curve Type Modeling Unit Root Test: Augmented Dickey Fuller Test
(Sample period: 1971Q1 - 2007Q1)

Critical value used in ADF test
With constant: 5 percent =-2.88; 1 percent=-3.48
With constant and trend: 5 percent=-3.44; 1 percent=-4.03
Null hypothesis: The variable has a unit root.

	Level				First difference			
	Trend	Lag length *p*	ADF statistics	Estimated γ-1	Trend	Lag length *p*	ADF statistics	Estimated γ-1
cpi	No	5	-4.53 **	-0.01	Yes	4	-3.97 *	-0.29
cpit	No	4	-4.52 **	-0.02	Yes	3	-4.18 *	-0.31
cpint	No	4	-4.33 **	-0.01	Yes	3	-3.72 *	-0.25
ulc	No	2	-4.33 **	-0.02	Yes	1	-5.21 **	-0.52
ip	No	1	-2.44	-0.03	No	0	-6.53 **	-0.45
ipe	No	1	-2.47	-0.03	No	0	-7.34 **	-0.54
openness	No	4	0.71	0.01	No	3	-3.48 **	-0.96
reer	No	1	-3.48	-0.05	No	0	-3.48 **	-0.74
un	No	3	-1.28	-0.01	No	2	-4.38 **	-0.52
stock price	No	1	-2.50	-0.02	No	0	-8.43 **	-0.66
estate stock price	No	0	-1.33	-0.02	No	0	-11.27 **	-0.95
gap	No	0	-2.52	-0.08	No	0	-11.23 **	-0.93

Notes:
1 For a variable *y*, the following regression equation is considered:

$$\Delta y_t = a_1 + a_2 \cdot t + (\gamma - 1) \cdot y_{t-1} + \sum_{i=1}^{p} \beta_i \Delta y_{t-i} + \varepsilon_t$$

where Δ is the difference operator, t a time trend, and *p* the lag length. The ADF statistic is the t ratio on the lagged variable, y_{t-1}

The sample period is, in principle, 1971Q1-2007Q1. However, for a model with a lag length 4 or 5, the sample period starts 1971Q2 or 1971Q3, respectively.

2. ** indicates significant at 1 percent level and * indicates significant at 5 percent level. The null hypothesis is that "the variable has a unit root": i.e., γ-1=0.

Table I.3. Phillips Curve Type Modeling Stationarity Test
— Kwiatkowski-Phillips-Schmidt-Shin (KPSS) Test
(Sample period: 1971Q1 - 2007Q1) 1/

Critical value used in KPSS test
With constant: 5 percent =0.46; 1 percent=0.74
With constant and trend: 5 percent=0.15; 1 percent=0.22
The null hypothesis: The variable is stationary.

	Level		First difference	
	Trend	KPSS statistics	Trend	KPSS statistics
cpi	Yes	0.32 **	Yes	0.16 *
cpit	Yes	0.31 **	Yes	0.12
cpint	Yes	0.33 **	Yes	0.15
ulc	Yes	0.31 **	Yes	0.15
ip	No	0.25	No	0.18
ipe	Yes	0.19 *	No	0.17
openness	Yes	0.30 **	No	0.11
reer	Yes	0.21 *	No	0.18
un	Yes	0.15 *	No	0.11
stock price	Yes	0.31 **	Yes	0.08
estate stock price	Yes	0.23 *	No	0.11
gap	No	0.09	No	0.04

1/ The sample period is, in principle, 1971Q1-2007Q1. However, for a model with a lag length 4 or 5, the sample period starts 1971Q2 or 1971Q3, respectively.

Table I.4. Traditional Phillips Curve Markup Type Inflation Models: VAR Lag Selections
(Sample period: 1972Q1 - 2007Q1)

Endogenous variables	cpi and ulc				cpi, ulc, and ip				cpi, ulc, and ipe			
	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test
Lag 8	-13.178	-13.724	-14.098	...	-15.425	-16.542	-17.307	...	-14.066	-15.183	-15.948	...
Lag 7	-13.290	-13.786	-14.126	0.842	-15.665	-16.670	-17.359	0.937	-14.288	-15.294	-15.982	1.157
Lag 6	-13.415	**-13.862**	**-14.168**	**0.459**	-15.920	-16.814	**-17.426**	**0.760**	-14.549	-15.443	**-16.055**	**0.689**
Lag 5	**-13.439**	-13.836	-14.108	3.641 **	-16.082	**-16.864**	-17.399	2.037 *	-14.716	**-15.498**	-16.033	1.972 *
Lag 4	-13.328	-13.676	-13.914	8.296 **	-16.064	-16.734	-17.193	4.695 **	-14.699	-15.369	-15.828	4.681 **
Lag 3	-13.367	-13.665	-13.869	3.290 *	-16.200	-16.759	-17.141	2.508 **	-14.861	-15.420	-15.802	2.130 *
Lag 2	-13.338	-13.586	-13.756	5.672 **	**-16.300**	-16.747	-17.053	3.107 **	**-14.942**	-15.389	-15.695	3.396 **
Lag 1	-12.935	-13.134	-13.270	20.255 **	-15.734	-16.069	-16.299	15.011 **	-14.544	-14.879	-15.109	11.716 **

Note:

1. For the sample period 1972Q1:2007Q1, the following VAR is estimated:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + B\ x_t + \varepsilon_t$$

where
y=the 2×1 or 3×1vector of endogenous variables.
x= the 7×1 vector of deterministic (exogenous) variables.
ε =the 2×1 or 3×1 vector of error terms.

2. Exogenous variables included in VAR models are constant, trend, and dummy variables for: the first oil shock (1974Q2=1), 1989 tax reform (1989Q2=1), 1997 tax reform (1997Q2=1), high growth periods (1972Q1-1973Q4=1), and zero interest rate policy periods (1999Q1-2006Q1=1).

3. SC, Schwarz criterion; HQ, Hannan-Quinn criterion; and AIC, Akaike information criterion. The bold figures indicates the smallest values of the information criteria.

4. The null hypothesis for F test is that lag reduction is acceptable: e.g., for a model with 7 lags, the null hypothesis is that all the lags at 8th length are zero.

Table I.5. Traditional Models: VAR Model Diagnostic Tests

Full sample periods (1972Q1-2007Q1)

Endogenous variables	cpi and ulc		cpi, ulc, and ip		cpi, ulc, and ipe	
Lag order	6	5	6	2	6	2
Autocorrelation (AR1-5)	2.419 **	2.685 **	1.927 **	1.937 **	2.205 **	2.401 **
Normality	70.875 **	74.129 **	93.840 **	127.090 **	113.670 **	177.350 **
Heteroscedasticity	5.131 **	6.408 **	1.720 **	4.765 **	1.662 **	5.347 **

Excluding oil shock periods (1981Q1-2007Q1)

Endogenous variables	cpi and ulc		cpi, ulc, and ip		cpi, ulc, and ipe	
Lag order	4	3	4	2	4	2
Autocorrelation (AR1-5)	1.060	1.299	0.926	1.367	0.893	1.506 *
Normality	1.869	1.413	20.551 **	19.898 **	42.253 **	41.591 **
Heteroscedasticity	0.743	0.857	0.807	1.064	0.626	0.757

Table I.6. Traditional Case—Phillips Curve Markup Type Inflation Models: Johansen Cointegration Tests
(Sample period, 1971Q3-2007Q1)

	Unrestricted cointegration		
	Model 1	Model 2	Model 3
Normalized β cointegrating vector			
cpi	1.000	1.000	1.000
	...	...	...
ulc	-0.457 **	-0.385 **	-0.548 **
	0.052	*0.057*	*0.043*
ip	0.042 *	...	...
	0.020	...	...
ipe	...	0.029 **	...
	...	*0.012*	...
Trend	-0.002 **	-0.002 **	-0.002 **
	0.000	*0.000*	*0.000*
Speed of adjustment (α) parameters			
cpi equation	-0.075 **	-0.070 **	-0.097 **
	0.013	*0.012*	*0.017*
ulc equation	0.093 **	0.072 **	0.096 **
	0.029	*0.027*	*0.039*
ip equation	-0.478 **	-0.818 **	...
	0.143	*0.255*	...
LR tests: α coefficients are zero	...	...	...
Trace tests			
r=0	66.568 **	69.339 **	46.526 **
	[0.000]	[0.000]	[0.000]
r≤1	8.336	8.342	0.605
	[0.976]	[0.976]	[1.000]
r≤2	1.058	0.493	...
	[0.999]	[1.000]	...
Other settings			
Unrestricted variables	Constant, oil74, tax89, tax97, high growth, and zero interest		
VAR lag order	6	6	6

Notes:
1. *Italic* indicates standard errors; [ ] indicates p value; and bold indicates parameter restrictions.
2. ** indicates significant at 1 percent level and * significant at 5 percent level.

Table I.7. Tradable Sector Inflation Models: Johansen Cointegration Tests
(Sample period, 1971Q3-2007Q1)

	Unrestricted cointegration					Model 4:Restricted cointegration	
	Model 1	Model 2	Model 3	**Model 4**	Model 5	Model 4a	Model 4b
Normalized β cointegrating vector							
cpit	1.000	1.000	1.000	1.000	1.000	1.000	1.000
	...	...	...	...	...	...	...
ulc	-0.596 **	-0.569 **	-0.539 **	-0.533 **	-0.554 **	-0.522 **	-0.543 **
	0.051	*0.052*	*0.048*	*0.032*	*0.048*	*0.035*	*0.037*
ip	...	0.034	...	...	...	...	...
	...	*0.020*	...	...	...	...	...
ipe	...	...	0.022 *	...	...	...	...
	...	...	*0.010*	...	...	...	...
openness	...	...	...	*0.329* **	...	0.341 **	0.285 **
	...	...	...	*0.049*	...	*0.052*	*0.056*
reer	...	...	...	...	-0.084 *	...	...
	...	...	...	...	*0.035*	...	...
Trend	-0.001 **	-0.001	-0.001 **	-0.003 **	-0.001 **	-0.003 **	-0.003 **
	0.000	*0.000*	*0.000*	*0.000*	0.000	*0.000*	*0.000*
Speed of adjustment (α) parameters							
cpit equation	-0.110 **	-0.085 **	-0.099 **	-0.163 **	-0.106 **	-0.178 **	-0.168 **
	0.027	*0.024*	*0.023*	*0.031*	*0.027*	*0.029*	*0.032*
ulc equation	0.064	0.094 *	0.081 *	0.127 *	0.070	**0.000**	0.153 **
	0.048	*0.040*	*0.040*	*0.057*	*0.048*	*...*	*0.059*
ip equation	...	-0.634	...	...	...	...	...
	...	*0.189*	...	...	...	...	...
ipe equation	...	...	-1.233 **	...	...	...	...
	...	...	*0.368*	...	...	...	...
openness equation	...	...	...	*-0.433* **	...	-0.441 **	**0.000**
	...	...	...	*0.102*	...	*0.099*	*...*
reer equation	...	...	...	...	0.436	...	...
	...	...	...	...	*0.182*	...	...
LR tests: α coefficients are zero	...		...	...	...	5.437 *	18.889 **
						[0.020]	[0.000]
Trace tests							
r=0	26.167 *	38.724 *	48.742 *	67.929 **	46.302 *	...	...
	[0.046]	[0.019]	[0.012]	[0.000]	[0.022]	...	...
r≤1	3.365	8.074	10.507	7.042	14.262	...	...
	[0.8307]	[0.680]	[0.902]	[0.993]	[0.636]	...	...
r≤2	...	0.077	3.8318	2.323	4.602	...	...
	...	[0.781]	[0.766]	[0.946]	[0.654]	...	...
Other settings							
Unrestricted variables	Constant, oil74, tax89, tax97, high growth, and zero interest						
VAR lag order	6	5	6	6	5	6	6

Notes:
1. *Italic* indicates standard errors; [ ] indicates p value; and bold indicates parameter restrictions.
2. ** indicates significant at 1 percent level and * significant at 5 percent level.

Table I.8. Non-tradable Sector Inflation Models: Johansen Cointegration Tests
(Sample period, 1971Q3-2007Q1)

	Unrestricted cointegration						Restricted cointegration	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	**Model 4**	**Model 5**
Normalized β cointegrating vector								
cpint	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
	...	...	...	...	...	...	...	...
ulc	-0.600 **	-0.557 **	-0.458 **	-0.700 **	-0.548 **	-0.577 **	-0.710 **	-0.561 **
	0.040	*0.046*	*0.060*	*0.037*	*0.050*	*0.043*	*0.036*	*0.048*
ip	...	0.050 **	...	...	...	...	...	...
	...	*0.018*	...	...	...	...	...	...
ipe	...	...	0.031 *	...	...	...	...	...
	...	...	*0.012*	...	...	...	...	...
un	...	...	...	0.013 *	...	...	0.012 *	...
	...	...	...	*0.005*	...	...	*0.005*	...
stock price	...	...	...	...	-0.024 *	...	...	-0.026 **
	...	...	...	...	*0.010*	...	...	*0.009*
estate stock price	...	...	...	...	...	-0.014		
	...	...	...	...	...	*0.008*		
Trend	-0.003 **	-0.002	-0.003 **	-0.003 **	-0.003 **	-0.002 **	-0.003 **	-0.003 **
	0.000	*0.000*	*0.000*	*0.000*	*0.000*	*0.000*	*0.000*	*0.000*
Speed of adjustment (α) parameters								
cpint equation	-0.093 **	-0.077 **	-0.065 **	-0.144 **	-0.096 **	-0.088 **	-0.149 **	-0.103 **
	0.012	*0.010*	*0.008*	*0.018*	*0.012*	*0.010*	*0.017*	*0.012*
ulc equation	0.046	0.046	0.017	0.034	0.019	0.015	**0.000**	**0.000**
	0.040	*0.035*	*0.027*	*0.062*	*0.040*	*0.035*	...	...
ip equation	...	-0.586 **	...	...	...	...	...	...
	...	*0.138*	...	...	...	...	...	...
ipe equation	...	...	-0.756 **	...	...	...	...	...
	...	...	*0.217*	...	...	...	...	...
un equation	...	...	...	0.382	...	...	**0.000**	...
	...	...	...	*0.574*	...	...	...	...
stock price equation	...	...	...	...	-0.419	...	...	**0.000**
	...	...	...	...	*0.267*	...	...	...
estate stock price equation	...	...	...	...	...	-0.326		
	...	...	...	...	...	*0.425*		
LR tests: α coefficients are zero	...		...	...	...	...	0.843	2.603
							[0.656]	[0.272]
Trace tests								
r=0	65.059 **	42.915 **	83.122 **	79.042 **	80.372 **	79.802 **	...	...
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	...	...
r≤1	1.368	25.872	13.088	11.993	12.994	10.361	...	...
	[0.9943]	[0.948]	[0.731]	[0.813]	[0.739]	[0.909]	...	...
r≤2	...	12.518	0.9811	2.585	4.0377	2.099	...	...
	...	[0.978]	[0.999]	[0.912]	[0.737]	[0.962]	...	...
Other settings								
Unrestricted variables	Constant, oil74, tax89, tax97, high growth, and zero interest							
VAR lag order	6	6	5	6	6	5	6	6

Notes:
1. *Italic* indicates standard errors; [ ] indicates p value; and bold indicates parameter restrictions.
2. ** indicates significant at 1 percent level and * significant at 5 percent level.

Table I.9. Augmented Aggregate Philipps-curve Type Inflation Models: Johansen Cointegration Tests
(Sample period, 1971Q3-2007Q1)

	Unrestricted cointegration		Restricted cointegration
	Model 1	Model 2	**Model 1**
Normalized β cointegrating vector			
cpi	1.000	1.000	1.000
	...	...	...
ulc	-0.571 **	-0.438 **	-0.543 **
	0.040	*0.056*	*0.044*
openness	0.202 **	0.296 **	0.224 **
	0.048	*0.095*	*0.052*
un	0.015 *	...	0.018 **
	0.005	...	*0.006*
stock price	...	-0.007	...
	...	*0.012*	...
Trend	-0.004 **	-0.004	-0.004 **
	0.000	*0.001*	*0.000*
Speed of adjustment (α) parameters			
cpi equation	-0.152 **	-0.079 **	-0.149 **
	0.025	*0.016*	*0.022*
ulc equation	0.078	0.088 *	**0.000**
	0.059	*0.036*	*...*
openness equation	-0.417 **	-0.219 **	-0.412 **
	0.109	*0.067*	*0.098*
un equation	-0.298	...	**0.000**
	0.583	...	*...*
stock price equation	...	-0.773 **	...
	...	*0.260*	...
LR tests: α coefficients are zero	...	...	2.233
			[0.327]
Trace tests			
r=0	101.650 **	92.614 **	...
	[0.000]	[0.000]	...
r≤1	38.414	37.769	...
	[0.131]	[0.149]	...
r≤2	14.050	...	...
	[0.654]	...	...
Other settings			
Unrestricted variables	Constant, oil74, tax89, tax97, high growth, and zero interest		
VAR lag order	6	5	6

Notes:
1. *Italic* indicates standard errors; [ ] indicates p value; and bold indicates parameter restrictions.
2. ** indicates significant at 1 percent level and * significant at 5 percent level.

Table I.10. Phillips Curve Type Inflation Functions:
Encompassing Tests Summary 1/

		Alternative models		
		Model 1	Model 2	Model 3
Null models	Model 1	...	*-1.033* [0.304]	*-1.842* [0.068]
	Model 2	*2.842* [0.005]	...	*0.272* [0.786]
	Model 3	*3.634* [0.000]	*1.597* [0.113]	...

1. *Italic* indicates t statistics; and [ ] indicates p value. High P value indicates that the fitted value of the alternative model is not significant, and therefore, the null model encompasses the alternative model. For example, figures in column 2 and row 1 show that if the fitted value of model 2 is included in model 1, its coefficient is insignificant with the t value of -1.033 and the probability value of 0.304.

Table I.11. Monetary Type Modeling: Descriptive Statistics
(In logarithm, 1970Q1 - 2007Q1)

	Mean	Standard deviation	Skewness
Level			
rm2cd	8.268	0.457	-0.370
rgdp	11.437	0.316	-0.516
oc	2.876	1.174	-0.006
real stock price	-0.247	0.451	0.459
real estate stock price	2.160	0.571	0.497
First difference			
Drm2cd	0.011	0.013	-0.360
Drgdp	0.007	0.009	-0.892
Doc	-0.016	0.444	-0.122
Dreal stock price	0.005	0.080	-0.151
Dreal estate stock price	0.013	0.143	0.322

Table I.12. Monetary Type Modeling Unit Root Test: Augmented Dickey Fuller Test
(Sample period: 1971Q1 - 2007Q1)

Critical value used in ADF test
With constant: 5 percent =-2.88; 1 percent=-3.48
With constant and trend: 5 percent=-3.44; 1 percent=-4.03
Null hypothesis: The variable has a unit root.

	Level				First difference			
	Trend	Lag length *p*	ADF statistics	Estimated γ-1	Trend	Lag length *p*	ADF statistics	Estimated γ-1
rm2cd	No	4	-1.80	0.00	No	1	-3.46 *	-0.20
	Yes	2	-2.79	-0.02	Yes	1	-3.46 *	-0.22
rgdp	No	0	-4.60 **	-0.01	No	2	-4.12 **	-0.44
	Yes	0	-1.20	-0.01	Yes	0	-9.97 **	-0.82
oc	No	0	-2.05	-0.06	No	1	-11.80 **	-1.45
	Yes	0	-4.15 *	-0.20	Yes	1	-11.79 **	-1.46
real stock price	No	1	-1.85	-0.03	No	0	-8.41 **	-0.65
	Yes	1	-1.81	0.70	Yes	0	-8.40 **	-0.65
real estate stock price	No	0	-0.66	-0.01	No	0	-11.14 **	-0.92
	Yes	0	-1.51	-0.04	Yes	0	-11.14 **	-0.93

Notes:
1 For a variable *y*, the following regression equation is considered:

$$\Delta y_t = a_1 + a_2 \cdot t + (\gamma - 1) \cdot y_{t-1} + \sum_{i=1}^{p} \beta_i \Delta y_{t-i} + \varepsilon_t$$

where Δ is the difference operator, t a time trend, and *p* the lag length. The ADF statistic is the t ratio on the lagged variable, y_{t-1}

2. ** indicates significant at 1 percent level and * indicates significant at 5 percent level. The null hypothesis is that "the variable has a unit root": i.e., γ-1=0.

Table I.13. Monetary Inflation Models: Johansen Cointegration Tests
(Sample period, 1971Q3-2007Q1)

	Unrestricted cointegration			Restricted cointegration
	Model 1	Model 2	Model 3	**Model 2**
Normalized β cointegrating vector				
rm2cd	1.000	1.000	1.000	1.000
	...	...	...	...
rgdp	-1.234 **	-1.280 **	-1.260 **	-1.277 **
	0.049	*0.034*	*0.052*	*0.034*
oc	0.081 **	0.035 **	0.052 **	0.034 **
	0.011	*0.008*	*0.010*	*0.008*
real estate stock price	...	-0.040 **	...	-0.044 **
	...	*0.013*	...	*0.013*
real stock price	...	...	-0.032	...
	...	...	*0.021*	...
Trend	...	...	...	...
	...	...	...	...
Speed of adjustment (α) parameters				
rm2cd equation	-0.047 **	-0.085 **	-0.060 **	-0.086 **
	0.016	*0.021*	*0.017*	*0.021*
rgdp equation	0.030 *	0.043 *	0.034 *	0.045 *
	0.016	*0.021*	*0.016*	*0.021*
oc equation	-3.678 **	-4.413 **	-4.071 **	-4.630 **
	0.903	*1.257*	*1.011*	*1.269*
real estate stock price	...	-0.472	...	**0.000**
	...	*0.422*	...	...
real stock price equation	...	...	-0.073	...
	...	...	*0.178*	...
LR tests: α coefficients are zero	...		...	1.254
				[0.263]
Trace tests				
r=0	45.651 **	63.574 *	59.102 *	...
	[0.000]	[0.001]	[0.003]	...
r≤1	9.561	18.753	19.952	...
	[0.316]	[0.511]	[0.426]	...
r≤2	3.842	4.377	6.6199	...
	[0.311]	[0.871]	[0.622]	...
Other settings				
Unrestricted variables	Constant, oil74, tax89, tax97, high growth, and zero interest			
VAR lag order	4	2	2	2

Notes:
1. *Italic* indicates t statistics; [ ] indicates p value; and bold indicates parameter restrictions.
2. ** indicates significant at 1 percent level and * significant at 5 percent level.

Table I.14. Phillips Curve Markup Type Versus Monetary Type Inflation Functions: Encompassing Tests Summary 1/

		Alternative models	
		Model 1	Model 4
Null models	Model 1	...	*2.808* [0.006]
	Model 4	*5.916* [0.000]	...

1. *Italic* indicates t statistics; and [ ] indicates p value. High P value that the fitted value of the alternative model is not significant, and therefore, the null model encompasses the alternative model.

Appendix Table I.1. Tradable Sector Inflation Models: VAR Lag Selections
(Sample period: 1972Q1 - 2007Q1)

Endogenous variables	cpit and ulc (Model 1)				cpit, ulc, and ip (Model 2)				cpit, ulc, and ipe (Model 3)				cpit, ulc, and openness (Model 4)				cpit, ulc, and reer (Model 5)			
	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test
Lag 8	-12.605	-13.176	-13.567	...	-14.818	-15.973	-16.763	...	-13.389	-14.543	-15.334	...	-16.837	-17.992	-18.782	...	-15.018	-16.173	-16.963	...
Lag 7	-12.726	-13.247	-13.604	0.559	-15.036	-16.079	-16.793	1.204	-13.604	-14.647	-15.361	1.231	-17.035	-18.078	**-18.791**	1.455	-15.240	-16.283	-16.997	1.148
Lag 6	-12.823	**-13.295**	**-13.618**	1.311	-15.276	-16.207	-16.845	0.950	-13.856	-14.788	**-15.425**	0.798	-17.204	-18.135	-18.773	1.864	-15.478	-16.409	-17.046	0.983
Lag 5	-12.859	-13.281	-13.570	3.240 *	-15.471	**-16.291**	**-16.851**	1.568	-14.037	-14.856	-15.417	1.763	-17.343	**-18.162**	-18.723	2.333 *	-15.691	-16.511	**-17.071**	1.323
Lag 4	-12.815	-13.187	-13.442	5.937 **	-15.505	-16.213	-16.697	3.891 **	-14.102	-14.810	-15.294	3.439 **	-17.411	-18.119	-18.603	3.388 **	-15.737	-16.445	-16.929	3.716 **
Lag 3	-12.877	-13.200	-13.421	2.496 *	-15.656	-16.252	-16.660	2.274 *	-14.279	-14.875	-15.283	1.904	-17.490	-18.086	-18.493	3.325 **	-15.894	-16.490	-16.898	2.187 *
Lag 2	**-12.902**	-13.175	-13.362	3.785 **	**-15.792**	-16.276	-16.608	2.554 **	**-14.422**	**-14.906**	-15.238	2.448 *	**-17.534**	-18.018	-18.349	3.930 **	**-16.047**	**-16.531**	-16.862	2.309 *
Lag 1	-12.289	-12.512	-12.665	29.513 **	-14.985	-15.357	-15.612	19.985 **	-13.792	-14.165	-14.420	16.196 **	-16.981	-17.353	-17.608	14.628 **	-15.501	-15.874	-16.129	14.475 **

Note:
1. For the sample period 1972Q1:2007Q1, the following VAR is estimated:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + B\ x_t + \varepsilon_t$$

where
y=the 2×1 or 3×1vector of endogenous variables.
x= the 7×1 vector of deterministic (exogenous) variables.
ε =the 2×1 or 3×1 vector of error terms.

2. Exogenous variables included in VAR models are constant, trend, and dummy variables for: the first oil shock (1974Q2=1), 1989 tax reform (1989Q2=1), 1997 tax reform (1997Q2=1), high growth periods (1972Q1-1973Q4=1), and zero interest rate policy periods (1999Q1-2006Q1=1).
3. SC, Schwarz criterion; HQ, Hannan-Quinn criterion; and AIC, Akaike information criterion. The bold figures indicates the smallest values of the information criteria.
4. The null hypothesis for F test is that lag reduction is acceptable: e.g., for a model with 7 lags, the null hypothesis is that all the lags at 8th length are zero.

Appendix Table I.2. Non Tradable Sector Inflation Models: VAR Lag Selections
(Sample period: 1972Q1 - 2007Q1)

Endogenous variables	cpint and ulc (Model 1)				cpint, ulc, and ip (Model 2)				cpint, ulc, and ipe (Model 3)				cpint, ulc, and un (Model 4)				cpint, ulc, and stock price (Model 5)				cpint, ulc, and estate stock price (Model 6)			
	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test
Lag 8	-13.399	-13.970	-14.361	...	-15.681	-16.835	-17.625	...	-14.153	-15.308	-16.098	...	-13.921	-15.075	-15.866	...	-14.765	-15.919	-16.709	...	-13.329	-14.484	-15.274	...
Lag 7	-13.514	-14.035	-14.392	0.739	-15.904	-16.947	**-17.661**	1.132	-14.364	-15.407	-16.121	1.289	-14.126	-15.169	-15.883	1.361	-14.962	-16.005	-16.719	1.453	-13.521	-14.564	-15.278	1.527
Lag 6	-13.613	-14.085	**-14.408**	1.227	-16.060	-16.991	-17.628	2.041 *	-14.584	-15.515	-16.152	1.208	-14.316	-15.248	**-15.885**	1.587	-15.202	-16.133	-16.770	0.958	-13.770	-14.701	-15.338	0.840
Lag 5	-13.671	**-14.093**	-14.382	2.549 *	-16.254	-17.074	-17.635	1.577	-14.801	-15.620	**-16.181**	1.279	-14.501	-15.320	-15.881	1.711	-15.427	-16.247	**-16.807**	1.169	-13.971	-14.791	**-15.351**	1.486
Lag 4	-13.617	-13.990	-14.245	6.277 **	-16.402	-17.109	-17.594	2.268 *	-14.933	-15.641	-16.125	2.481 **	-14.592	-15.300	-15.784	3.062 **	-15.480	-16.187	-16.672	3.616 **	-14.027	-14.735	-15.219	3.569 **
Lag 3	**-13.750**	-14.073	-14.294	0.233	-16.608	-17.204	-17.612	1.499	-15.164	-15.760	-16.168	1.156	-14.822	-15.418	-15.826	1.168	-15.703	-16.299	-16.706	1.265	-14.263	-14.859	-15.267	1.084
Lag 2	-13.743	-14.017	-14.203	4.849 **	**-16.727**	**-17.211**	-17.543	2.799 **	**-15.279**	**-15.763**	-16.095	2.859 **	-14.989	-15.473	-15.804	2.101 *	-15.828	**-16.312**	-16.643	2.711 **	-14.421	-14.905	-15.236	2.234 *
Lag 1	-13.746	-13.970	-14.123	4.595 **	-16.485	-16.857	-17.112	8.792 **	-15.232	-15.605	-15.860	5.481 **	**-15.164**	**-15.536**	-15.791	2.031 *	**-15.935**	-16.307	-16.562	3.060 **	**-14.561**	**-14.933**	-15.188	2.554 **

Note:
1. For the sample period 1972Q1:2007Q1, the following VAR is estimated:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + B\ x_t + \varepsilon_t$$

where
y=the 2×1 or 3×1vector of endogenous variables.
x= the 7×1 vector of deterministic (exogenous) variables.
ε =the 2×1 or 3×1 vector of error terms.

2. Exogenous variables included in VAR models are constant, trend, and dummy variables for: the first oil shock (1974Q2=1), 1989 tax reform (1989Q2=1), 1997 tax reform (1997Q2=1), high growth periods (1972Q1-1973Q4=1), and zero interest rate policy periods (1999Q1-2006Q1=1).
3. SC, Schwarz criterion; HQ, Hannan-Quinn criterion; and AIC, Akaike information criterion. The bold figures indicates the smallest values of the information criteria.
4. The null hypothesis for F test is that lag reduction is acceptable: e.g., for a model with 7 lags, the null hypothesis is that all the lags at 8th length are zero.

Appendix Table I.3. Augmented Aggregate Phillips-curve Type Inflation Models: VAR Lag Selections
(Sample period: 1972Q1 - 2007Q1)

Endogenous variables	cpi, ulc, openness, and ulc (Model 1)				cpi, ulc, openness, and stock price (Model 2)			
	SC	HQ	AIC	F test	SC	HQ	AIC	F test
Lag 8	-17.889	-19.826	**-21.151**	...	-18.377	-20.314	-21.639	...
Lag 7	-18.218	-19.956	-21.146	1.498	-18.747	-20.485	-21.675	1.225
Lag 6	-18.500	-20.039	-21.093	1.890 *	-19.173	-20.713	**-21.767**	0.892
Lag 5	-18.728	-20.069	-20.986	2.362 **	-19.502	**-20.842**	-21.760	1.624
Lag 4	-18.908	-20.050	-20.832	2.826 **	-19.623	-20.765	-21.547	3.291 **
Lag 3	-19.135	**-20.078**	-20.724	2.543 **	-19.809	-20.752	-21.398	2.879 **
Lag 2	**-19.282**	-20.027	-20.537	3.304 **	-19.922	-20.667	-21.177	3.600 **
Lag 1	-19.169	-19.715	-20.089	5.820 **	**-19.959**	-20.506	-20.880	4.403 **

Note:

1. For the sample period 1972Q1:2007Q1, the following VAR is estimated:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + B x_t + \varepsilon_t$$

where
y=the 4×1vector of endogenous variables.
x= the 7×1 vector of deterministic (exogenous) variables.
ε =the 4×1 vector of error terms.

2. Exogenous variables included in VAR models are constant, trend, and dummy variables for: the first oil shock (1974Q2=1), 1989 tax reform (1989Q2=1), 1997 tax reform (1997Q2=1), high growth periods (1972Q1-1973Q4=1), and zero interest rate policy periods (1999Q1-2006Q1=1).

3. SC, Schwarz criterion; HQ, Hannan-Quinn criterion; and AIC, Akaike information criterion. The bold figures indicates the smallest values of the information criteria.

4. The null hypothesis for F test is that lag reduction is acceptable: e.g., for a model with 7 lags, the null hypothesis is that all the lags at 8th length are zero.

Appendix Table I.4. Monetary Inflation Models: VAR Lag Selections
(Sample period: 1972Q1 - 2007Q1)

Endogenous variables	rm2cd, rgdp, and oc (Model 1)				rm2cd, rgdp, oc, and real estate stock price (Model 2)				rm2cd, rgdp, oc, and real stock price (Model 3)			
	SC	HQ	AIC	F test	SC	HQ	AIC	F test	SC	HQ	AIC	F test
Lag 8	-10.755	-11.888	-12.664	...	-10.269	-12.170	-13.471	...	-11.681	-13.582	-14.883	...
Lag 7	-11.040	-12.064	-12.764	0.301	-10.762	-12.468	-13.636	0.3632	-12.156	-13.862	-15.030	0.4859
Lag 6	-11.305	-12.219	-12.844	0.571	-11.205	-12.716	-13.751	0.7238	-12.582	-14.093	-15.127	0.8402
Lag 5	-11.538	-12.342	-12.893	1.010	-11.523	-12.839	-13.740	1.6667	-12.993	-14.309	-15.210	0.9818
Lag 4	-11.732	**-12.427**	**-12.902**	1.579	-11.870	-12.992	-13.759	1.5037	-13.298	-14.420	-15.187	1.8261 *
Lag 3	-11.909	-12.494	-12.895	1.862	-12.222	-13.148	-13.782	1.5231	-13.632	-14.558	-15.192	1.6671
Lag 2	**-12.103**	-12.578	-12.903	1.669	**-12.637**	**-13.368**	**-13.869**	1.0513	**-14.040**	**-14.771**	**-15.271**	1.1121
Lag 1	-11.493	-11.859	-12.109	16.122 **	-12.369	-12.905	-13.272	7.4492 **	-13.599	-14.135	-14.502	9.3000 **

Note:

1. For the sample period 1972Q1:2007Q1, the following VAR is estimated:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + B x_t + \varepsilon_t$$

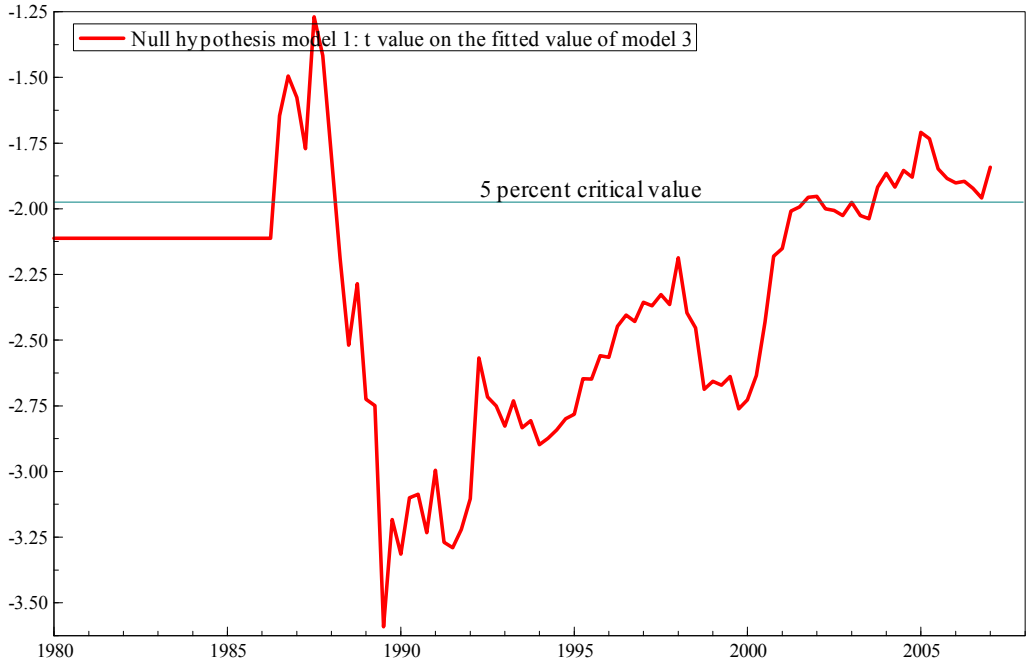
where
y=the 4×1vector of endogenous variables.
x= the 7×1 vector of deterministic (exogenous) variables.
ε =the 4×1 vector of error terms.

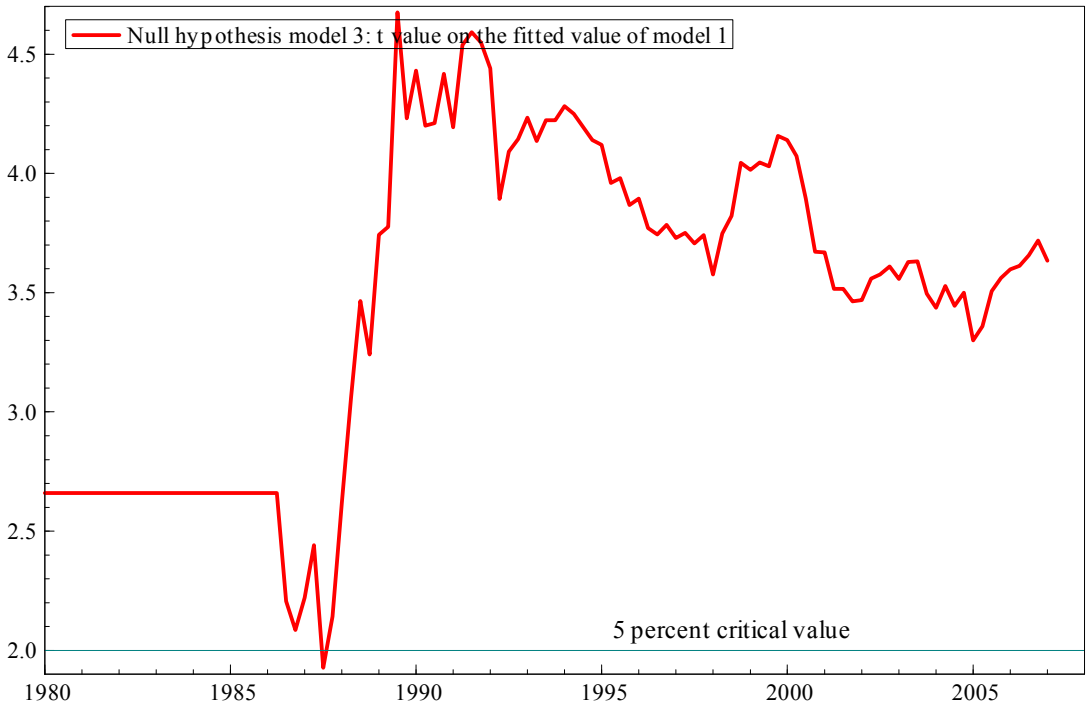
2. Exogenous variables included in VAR models are constant, trend, and dummy variables for: the first oil shock (1974Q2=1), 1989 tax reform (1989Q2=1), 1997 tax reform (1997Q2=1), high growth periods (1972Q1-1973Q4=1), and zero interest rate policy periods (1999Q1-2006Q1=1).

3. SC, Schwarz criterion; HQ, Hannan-Quinn criterion; and AIC, Akaike information criterion. The bold figures indicates the smallest values of the information criteria.

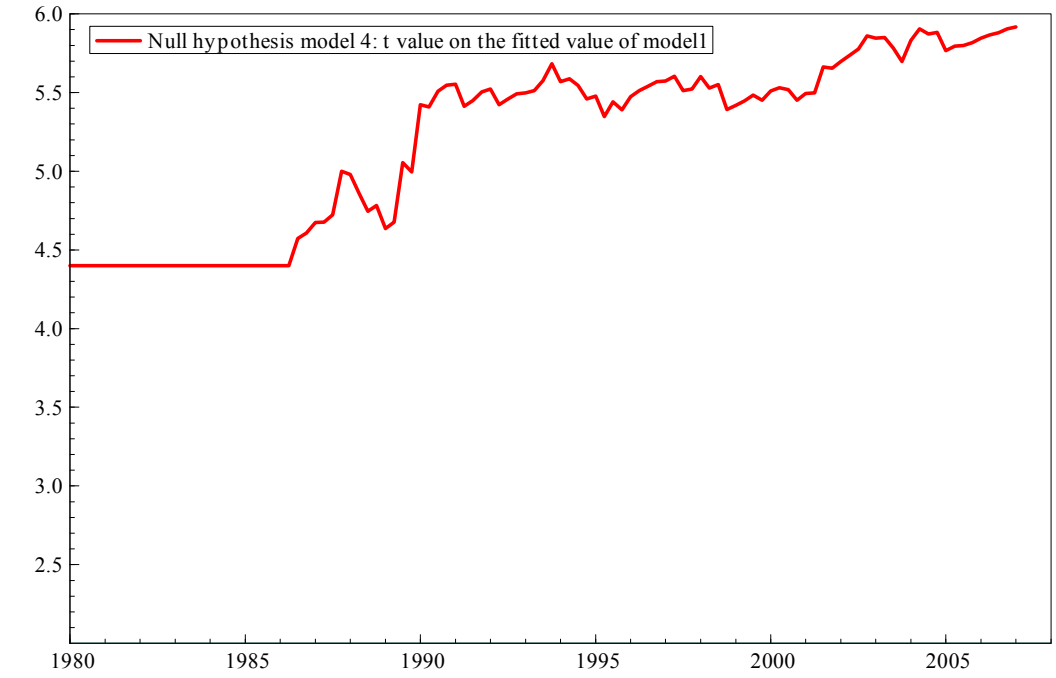
4. The null hypothesis for F test is that lag reduction is acceptable: e.g., for a model with 7 lags, the null hypothesis is that all the lags at 8th length are zero.

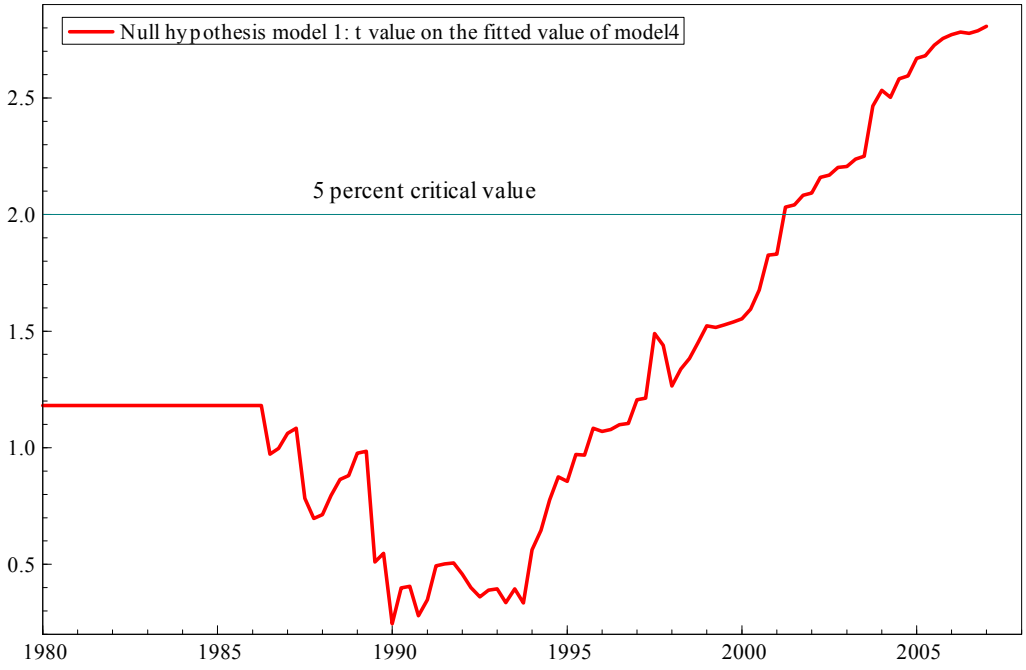
Appendix Figure I.1. Phillips Curve Type Inflation Functions Encompassing Tests: Recursive Estimates on *T* Values on β Coefficients on the Fitted Values





Appendix Figure I.2. Phillips Curve Type versus Monetary Type Inflation Functions Encompassing Tests:
Recursive Estimates on T Values on β Coefficients on the Fitted Values





Essay II: Money Demand, Dollarization, and High Inflation in Cambodia and Lao P.D.R.

I. INTRODUCTION

Cambodia and Lao P.D.R. have shared similar economic episodes over the last three decades. Both countries were devastated during the Indochina war periods in the 1960s and 1970s. Thereafter, both countries adopted a socialist command-type regime until the late 1980s. The failure of the command-type model became increasingly apparent by the middle of the 1980s, as per capita GDP growth lagged behind other economies in the region, inflation was much higher, and monetization remained significantly under-developed. The realization that the centralized planning approach could not provide incentives for economic growth, Cambodia and Lao P.D.R. changed their respective development strategy to gradually liberalize and open up the economy. In the past decade and half both countries have achieved remarkable economic developments, together with increased monetization of economy, albeit from a very low base.

An important common feature of Cambodia and Lao P.D.R. economy is their high level of dollarization. This poses a major challenge to policymakers, particularly in conduct of monetary policy. For example, the extensive use of foreign currencies (mainly U.S. dollars in Cambodia, and U.S. dollars and Thai baht in Lao P.D.R.) in economic activity weakens the central bank's control over money supply. Dollarization has never been officially endorsed by the monetary authorities in either of country. Rather the economy has been dollarized with the massive flows of foreign currencies occurring from aid by foreign

donors and later in private capital. Macroeconomic policy mismanagement (both countries experienced nearly hyperinflation and large exchange rates devaluations in the 1990s) and political uncertainties (particularly in the case of Cambodia) resulted in loss of confidence in the local currency and wider use of foreign currencies.

Nevertheless, an interesting difference between two countries exists. In Cambodia, the degree of dollarization, as measured by the ratio of foreign currency deposits to the total broad money, has continuously increased, while in Lao P.D.R., it has fluctuated and has even declined in recent years, implying a relative increase in the demand for local currency money (Figure II.1).

My research looks at the nature of money demand in both countries. I am interested in understanding whether dollarization is reversible and or sound economic policy alone could rebuild confidence in local currency. The comparison of Cambodia and Lao P.D.R.'s experiences provides an interesting experiment. There have been relatively few studies on both money demand and dollarization in the literature. Thus, my contribution is timely and may be useful.

My research addresses the question by analyzing money demand in Cambodia and Lao P.D.R., following an approach taken by Montiel (1989), Agénor and Khan (1996), Vetlov (2001), and Ericsson and Kamin (2003). If successfully found, a stable money demand function can assist the monetary authorities in Cambodia and Lao P.D.R. in guiding their monetary policy decisions. The structure of my paper is as follows. Section II provides a

literature review on money demand function and dollarization, and section III briefly overviews economic history in Cambodia and Lao P.D.R., necessary for understanding monetary and inflation developments in these countries. Section IV discusses data, and Section V presents empirical analysis of money demand in Cambodia and Lao P.D.R. Section VI presents some extension of the empirical model, and Section VII discusses policy implications. Section VII concludes.

II. LITERATURE REVIEW

There have been a number of studies of money demand in Asian developing countries. For example, Hafer and Kutan (1994) and Tseng (1994) estimated a money demand function for China, Moosa (1992) for India, Price and Insukindro (1994) and Dekle and Pradhan (1997) for Indonesia and Thailand, Arize (1994) for Korea, and Sriram (1999) for Malaysia. Sriram (2001) documented a comprehensive literature review of empirical money demand studies. Many of these studies were motivated by concerns about the stability of a money demand function amid rapidly changing circumstances, such as financial liberalization and increased openness of the economy, which happened in the late 1980s and 1990s.

The literature on dollarization, in turn, has developed more in the context of analyzing monetary economy in Latin America and Central and Eastern Europe. Dollarization and currency substitution are often used interchangeably in the literature, although many early

papers discussed the definition of the phenomenon. [1] An example of terminology often cited in the literature is the one defined by Cuddington (1989) and Calvo and Végh (1992). They refer to currency substitution as the use of foreign currency for a of medium of exchange. Dollarization implies a foreign currency is used as a store of value and unit of account. Calvo and Végh (1992) and Sahay and Végh (1995) introduce the concept of asset substitution. This applies to a situation where foreign-currency-denominated assets are primary used as a store of value. There is a large volume of empirical research about dollarization, investigating the determinants of the dollarization phenomenon: for Latin American countries, Ortiz (1983), Ramirez-Rojas (1985), Guidotti and Rodriguez (1991), Savastano (1992), and Ericsson and Kamin (2003); and for Central and Eastern European transition countries, Sahay and Végh (1995), Korhonen (1996), Brodsky (1997), Mongardini and Mueller (1999), Sarajevs (2000), Bichs (2000), and Vetlov (2001). One of the common findings of these studies is that dollarization occurred in reaction to the pressure of high inflation, including negative domestic real interest rates, and large devaluations of exchange rates. However, studies on dollarization in Asian countries have been rather limited, partly reflecting limited cases of the dollarization phenomenon.

As discussed in Sahay and Végh (1995), dollarization often follows high inflation. In this context, there are voluminous studies about hyper- and high inflations and money demand, with Cagan (1956)'s pioneer work in this area. Many of the recent empirical studies on hyper inflation employ the cointegration technique to estimate a money demand function

[1] In this paper, we do not particularly distinguish between dollarization and currency substitution, partly because it is not easy to empirically distinguish two concepts. It is not

under high inflation situations: for example, Taylor (1991), Phylakis and Taylor (1993), Engsted (1993, 1994, and 1998), Michael, Nobay, and Peel (1994), Choudhry (1998), Petrovic, Bogetic, and Vujosevic (1999), and Christev (2005). The main findings of these studies provide some empirical support for a Cagan-type money demand function for most hyper- or high inflation episodes. The only formal empirical study on monetary demand or dollarization in Cambodia and Lao P.D.R. that I am aware of is De Zamaróczy and Sa (2003). Their paper employs a cointegration technique to examine characteristics of money demand in highly dollarized and once-nearly hyperinflationary economies, Cambodia and Lao P.D.R.

III. Inflation and Monetary Developments in Cambodia and Lao P.D.R.

Following the civil war periods in the early 1970s, both Cambodia and Lao P.D.R. adopted a command-type central planning system. By comparison, Cambodia went far extreme, taking the extreme revolutionary experience of the Khmer Rouge during 1975-79, which involved bans on private property, as well as banking and money, leading to catastrophic economic disaster and human losses. The Khmer Rouge was driven out of most parts of Cambodia by the Vietnamese in 1979. Under the new transition regime, the local currency, riel, was reintroduced in 1980, while a state-owned mono bank was established at the same time, which played an important role in central, commercial, and development banking activities, according to the socialist economic model. Lao P.D.R. was founded in 1975 after the end of the Indochina war. The Lao-style central planning system can be in principle

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also focus of this paper.

characterized as a high decree of centralization of economic decision making and strict control and little reliance on the market mechanism of prices. Thus, what was produced, by whom, and for what uses were mainly subject to administrative decisions (Saignasith, 1997). Under this system, prices were controlled, domestic and external trade was strictly regulated, and state-owned enterprises and banks played a major role in providing goods and services but were strictly controlled by the government and without any room for managerial flexibility of autonomy.

In both countries, the failure of the command economy model became increasingly apparent by the middle of the 1980s, and the governments gradually started opening up the economy. During the 1980s, both countries achieved only limited monetization and most domestic transactions were based on barter. In Cambodia, the United Nations dispatched humanitarian and emergency aid and remittances from abroad resumed. However, with the disengagement of the Vietnamese in 1988-91, the government filled resulting fiscal gaps by recourse to central bank financing, resulting in very high inflation in the range of 90 to 180 percent a year.

In Lao P.D.R., the government launched a major economic reform, called "New Economic Mechanism," in 1989 to transform its economic system from a socialist-based to a marked-based economy. The reform program was implemented progressively and gradually, but covered all economic sectors. Reform in the state-owned enterprise sector led to a drop in budgetary transfers, while the new tax system remained rudimentary. Increases in civil service wages, in part to monetize fringe benefits, contributed to rising wage bill. These all

together led to a widening of the fiscal deficit. Accordingly, money growth was high, reaching nearly 90 percent by 1989, on account of monetization of budget deficit and rapid credit expansion to state-owned enterprises. Expansionary fiscal and monetary policy, together with a severe contraction in agriculture output due to a protracted drought, triggered reacceleration in inflation in the late 1980s.

In the early 1990s, Cambodia and Lao P.D.R. embarked on macroeconomic stabilization programs, with assistance from foreign donors, including the International Monetary Fund and World Bank. Monetary policy and fiscal policy were tightened. Inflation dropped to single-digit levels. However, in Cambodia, large-scale donor operations together with the large number of the return of Cambodian refugees and expatriates from abroad brought dollars or Thai baht, and as a result, the Cambodian economy was rapidly dollarized.

In the second half of the 1990s, macroeconomic performance again deteriorated in both countries, partly reflecting adverse impacts of the effects of the Asian crisis. However, domestic factors also played a certain role in worsened economic situations. In Cambodia, the political situation deteriorated to the point that factional fighting erupted in July 1997. Economic growth was slowing and compounded by the onset of the Asian crises. The government resorted to budget financing through monetization from the central bank. Inflation picked up to double digits, while the exchange rate depreciated nearly 30 percent in 1997-99, although its depreciation was smaller than those of other Asian currencies (Figure II.2). In Lao P.D.R., the government implemented ambitious public investment plans, such as irrigation projects to achieve rice self-sufficiency, despite domestic resource

constraints. An expansion in central bank financing for budget deficits caused money supply and inflation to skyrocket to triple digits, and the exchange rate to depreciate sharply (90 percent) in 1997-99 (Figure II.3).

To restore macroeconomic balances, the authorities in both Cambodia and Lao P.D.R. reintroduced a stabilization program in mid-1999. By restraining bank credit and tightening fiscal policy, both money growth and inflation rates fell to single digits by the middle of 2000, and the exchange rate was stabilized. Since then, macroeconomic stability was broadly maintained in both countries supported by prudent monetary and fiscal policy, and in recent years, inflation rates have fallen to historical lows with stable exchange rates. However, as mentioned earlier, there is one interesting difference between the two countries in recent macroeconomic performance, namely, trends of demand for local currency or dollarization. In Cambodia, despite macroeconomic stability, the trend in dollarization has not abated. By contrast, in Lao P.D.R. experiencing sustained macroeconomic stability, there are signs of a reversal in dollarization. The degree of dollarization has steadily declined since 1999. Below, I will empirically investigate the different characteristics of money demand and dollarization in these neighboring countries.

IV. DATA

All data are monthly and available from January 1995 to March 2007. They are available from IMF International Financial Statistics. Main variables of interest in this paper are

monetary aggregates in local currency, inflation, nominal exchange rates, and real GDP (descriptive statistics are reported in Table II.1). [2]

- Monetary aggregates in local currency (*LC*): *LC* is defined as local currency in circulation and demand and time deposits excluding foreign currency deposits. Real monetary aggregates are defined by deflating *LC* by the consumer price index (*CPI*) and are named, *RLC*. The base year is 2000. The series obtained from the IMF have been seasonally adjusted using the X12 ARIMA model.

- Inflation (*dcpi*): Defined as monthly growth of consumer price index (*CPI*, seasonally adjusted) in percent.

- Nominal effective exchange rate (*NEER*): A larger value indicates appreciation. A change in the exchange rate is defined as monthly growth of *NEER* and is named, *dneer*.

- Real GDP (*RGDP*): Neither data for monthly real GDP nor other alternatives representing the scale of activities are available. Hence, as proxy, monthly real GDP data are estimated by applying a quadratic interpolation method to annual real GDP.

[2] Capital letters refer to the level of variables, while lower case letters refer to the difference (for logarithm) or percent change of variables.

The Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) unit root tests are applied to examine the order of integration of these variables. The test for each variable starts with a very general specification with 12 lags, which should be considered as enough lag length for monthly data, and then the model is simplified by dropping irrelevant trend or lags. The log levels of real currency and GDP for both countries cannot reject the null hypothesis of a unit root with the ADF and PP tests. We can reject the null hypothesis that the first differences have a unit root using the PP test. The critical values for the change in Cambodian real GDP and Lao P.D.R change in real currency cannot reject the null of a unit root. This suggests that they may follow an I(2) process. However, the numerical values for the coefficients of the tests are 0.75 and 0.47 respectively. Numerically, these appear distant from unity. I conclude that the real currency and real GDP series are I(1).

Except in a single case, the monthly change in the CPI and nominal effective exchange rates appear to be stationary processes; their levels are I(1). The ADF test for the Lao P.D.R. inflation rate does not reject the null hypothesis of a unit root and the difference of the inflation rate is stationary. Once again, the numerical results suggest that the inflation rate may be stationary. However, the ADF and PP tests are for the full sample 1995-2007 and it is very clear from the graphical plots of the two series that they are not covariance stationary. Both processes are different in the 1996-2000 period; there may have been mean shifts since then. The ambiguity of the integration order for those variables of interest affect the choice of econometric method to estimate long-run relationships, i.e., Johansen's approach (1998) the most widely used cointegration technique — may not be appropriate because there is uncertainty whether variables are I(1) or I(0). As Rhabek and Mosconi

(1999) showed, while Johansen's vector error correction model approach can be used with a mixture of I(1) and I(0) variables, the likelihood testing procedure for the cointegrating rank can be sensitive to the presence of stationary variables (see further discussion in Section V-B.)

V. Long-run Money Demand Function

A. Model

Following the standard monetary theory, a long-run money demand function can be expressed as following.

$$(1)\ \frac{M}{P} = L(Y, OC);$$

where M is demand for money, P is the price level (CPI), Y is the level of activities (RGDP), and OC is the opportunity cost of holding money.

For monetary aggregates, as discussed earlier, local currency money (LC) is chosen. As the measure of opportunity costs, expected inflation rates and an expected change in exchange rates are considered. In developing countries where financial markets are underdeveloped and monetization is relatively low, expected inflation rates, rather than interest rates, can be regarded as a better proxy for the opportunity cost of holding money. This implies that people's choice is between holding money and real assets in their asset allocations, and if an expected inflation rate rises, they would prefer holding more real assets to money. Also, the expected change in the exchange rate is a useful alternative candidate for the opportunity cost, given the high degree of dollarization in Lao P.D.R. A fall in the

exchange rate (deprecation) would raise values of foreign currency money in terms of local currency money, and thereby reducing demand for LC.

Replacing *M* with *LC* (and deflating it by CPI, the left hand side of the equation is RLC), and in the log linear form,

(2) $rlc = c + \beta_1 \times rgdp + \beta_2 \times oc$;

where *c* is the constant term. We estimate the following two models:

$rlc = c + \beta_1 \times rgdp + \beta_2 \times dcpi$; $\beta_1 > 0, \beta_2 < 0$;

$rlc = c + \beta_1 \times rgdp + \beta_2 \times dneer$; $\beta_1 > 0, \beta_2 > 0$;

Note that the inflation rate (*dcpi*) and change in the exchange rate *(dneer)* are expressed as a monthly percent change of CPI and NEER, not in the first difference of the logarithm. This implies that β_2 is the semi-elasticity of real money balances with respect to the expected inflation rate or change in the exchange rate. The full elasticity can be expressed as $\beta_2 \times$ *expected inflation (or change in the exchange rate)*, implying that the elasticity varies with the expected change in the opportunity cost. Also, I estimated models that exclude real GDP, given that (i) as Cagan (1958) argued that in an economy experiencing very high inflation, level of activities may not matter in determining money demand; and (ii) reliable estimates for monthly GDP or economic activity data are not available for Cambodia and Lao P.D.R.

$rlc = c + \beta_1 \times dcpi$; $\beta_1 < 0$; and

$rlc = c + \beta_1 \times dneer$; $\beta_1 > 0$.

What variable should be chosen as a proxy for expected inflation rates and changes in the exchange rate can be controversial. However, this paper takes the simplest approach. Assuming rational expectations and perfect foresight, an inflation rate or change in the exchange rate for T+1 period is used as a proxy for expectation at time T.

B. Modeling strategy

I interpret the above Cagan-type inflation model as describing an equilibrium relationship among the variables of interest (e.g., *rlc*, and *dcpi* or *dneer*). The empirical model shows a steady state relationship among these variables, given that these variables are integrated processes. In other words, if these variables are linked to form an equilibrium relationship in the long-run, they will move closely together over time. Such an interpretation is consistent with the objectives of my research because my main interest lies in the behavior of real local currency demand and its interactions with other relevant variables (e.g., *dcpi* or *dneer*) in the long-run. In this context, my research attempts to test whether any equilibrium relationship exists; and if so, whether it is stable.

Accordingly, the cointegration approach, suggested by Johansen (1988 and 1991), would be most suited for my research objectives. Among alternative cointegration techniques, Johansen's VAR approach can be considered as the most appropriate one, particularly

because this allows me to build a model without imposing any assumptions about endogeneity issues and to employ a general-to-specific approach.

However, there is uncertainty about whether the regressors are I(0) or I(1), and therefore, Johansen's approach to test a cointegration relationship may not be appropriate. Accordingly, I use the bounds test procedures for the analysis of long-run relationships within an auto-regressive distributed lag (ARDL) framework, developed by Pesaran and Shin (1999) and Pesaran et al. (2001). There are several advantages for the ARDL bounds test over the Johansen approach. First, the bound test approach does not require all of regressors to be integrated of the same order. As discussed earlier, ADF and PP unit root tests resulted in inconsistent conclusions, which may partly reflect the volatile nature of macroeconomic indicators for these developing economies where structural breaks are common. The ARDL bounds test would be suited because it can be applied for a mixture of I(1) and I(0) variables or fractionally integrated variables. Second, it is known that the ARDL bounds test is more robust for small samples than the Johansen approach (Ghatak and Siddiki, 2001).

This said, I fully recognize a disadvantage of the ARDL bound test approach. This approach does not allow me to answer a question about the endogeneity issue. Thus, issues related to causality—e.g., whether changes in inflation cause real local currency demand or vice versa—are beyond the scope of this paper.

C. Estimation

To apply the ARDL bounds test, I model the equation (2) as an ARDL form:

$$(3)\ \Delta rlc_t = c + \beta_1 \times rlc_{t-1} + \beta_2 \times rgdp_{t-1} + \beta_3 \times oc_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta rlc_{t-i}$$

$$+ \sum_{j=0}^{q} \gamma_j \Delta rgdp_{t-j} + \sum_{k=0}^{r} \gamma_k \Delta oc_{t-k} + \varepsilon_t$$

where c is the intercept term and ε_t are white noise errors. Note that $rgdp_{t-1}$ and $\sum_{j=0}^{q} \gamma_j \Delta rgdp_{t-j}$ can be dropped from the estimation model when we analyze the long run relationship only between *rlc* and *dcpi* or *dneer*.

According to Pesaran et al. (2001), if there is no long-run relationship among real local currency money, real GDP, and opportunity costs, the null hypothesis, $\beta_1 = \beta_2 = \beta_3 = 0$ would hold. Under the alternative hypothesis ($\beta_1 \neq 0$, $\beta_2 \neq 0$, and $\beta_3 \neq 0$), a stable long-run relationship would exist among these variables. If so, a long-run model can be obtained from (3), when $\Delta rlc = \Delta rgdp = \Delta oc = 0$,

$$(4)\ rlc_t = -\frac{c}{\beta_1} - \frac{\beta_2}{\beta_1} \times rgdp_t - \frac{\beta_3}{\beta_1} \times oc_t + \nu_t$$

$$= \theta_0 - \theta_1 \times rgdp_t - \theta_2 \times oc_t + \nu_t$$

where ν_t is a mean zero stationary process. Using Pesaran et al.'s terminology, *rgdp* and *oc* are the long-run forcing variables for real money demand in local currency. Expected sign conditions in the equation (4) are: $\theta_1 < 0 \Leftrightarrow \beta_1 and\ \beta_2$ have different signs;

$\theta_2 > 0 \Leftrightarrow \beta_1 and\ \beta_3$ have the same signs (when *oc=dcpi*) or $\beta_1 and\ \beta_3$ have different signs (when *oc=dneer*).

The equation (3) can be estimated by OLS. Then I use the table of F or Wald test statistics developed by Pesaran et al. (2001), for the bounds test assessing whether the null hypothesis is accepted. If the test statistics exceed the upper critical value, we can conclude that the null hypothesis of no long run relationship is rejected, and thus a long-run relationship exists. If the test statistics fall below the lower critical value, we cannot reject the null hypothesis of no long-run relationship. If the statistics fall between the upper and lower critical values, the result is inconclusive. These conclusions can be derived regardless whether the integration orders of the regressors are I(0) or I(1).

Prior to carrying out the bounds test, I have to determine the lag order of the equation (3). Table II.3 summarizes the results of the lag selection for ARDL models based on the F statistics (lag exclusion test), the Akaike Information Criterion (AIC), and Schwarz Bayesian Criterion (SC). Results from the three tests vary widely and therefore the lag selection is inconclusive. Thus, I consider two other criterion which assess whether ARDL estimated residuals have no serial correlation and are normally distributed. I began with the longest lags suggested by the above three tests, and analyze diagnostic statistics for the ARDL models, including error autocorrelation, normality, and heteroskedasticity. If the error terms satisfy these diagnostic test, I reduce the model by one lag and then analyze diagnostic statistics again; otherwise I maintain the longest lag order suggested by either F

statistics, AIC, or SC (in all cases, F statistics or AIC suggest the longest lag order). I select the shortest possible lag order which passes the diagnostic tests. The test results are generally satisfactory for all models for Lao P.D.R., except that Model 5. There appears to be evidence of heteroskedasticity (Table II.4). All models for Cambodia suffer from either no evidence of normality or heteroskedasticity, which could raise concern over whether we can obtain a cointegrating vector interpretable as demand for money relationship.

Table II.5 reports the results of ARDL bounds test (Wald test statistics). For Cambodia, the test statistics for models 1 and 3 fall below the lower critical value, suggesting no statistical evidence of cointegration. The test statistics would not reject the null hypothesis of no long-run relationship for models 2 and 4, but the signs for *dcpi* (model 2) and *dneer* (model 4) are not consistent with the expectations. By contrast, for Lao P.D.R., long-run relationships were found for models with *dcpi* or *dneer* as a long-run forcing variable (models 5 and 7). The test statistics for models 5 and 7 exceeded the upper critical value, with their signs for all coefficients meeting expectations. However, the test statistics for models 6 and 8 (both including *dgdp* additionally as a long-run forcing variable) fell between the upper and lower critical values, suggesting that we cannot conclude the existence of a long-run relationship.

Those results indicate that in Cambodia, demand for local money is not responsive to inflation or the change in exchange rate, while in Lao P.D.R., statistical evidence is confirmed that lower inflation or exchange rate appreciation is associated with an increase in local currency money demand. They further imply that monetary policy affecting local

monetary conditions may not be effective in Cambodia but may have some influence in Lao P.D.R. This conclusion is consistent with observations that despite recent continuous declines in inflation and appreciation of exchange rate, dollarization has continued in Cambodia, while de-dollarization has advanced in Lao P.D.R.

Various parameter constancy tests suggest a broad stability of the estimated models 5 and 7 for Lao P.D.R. For model 5, long-run recursive coefficients estimated forwardly suggest that all coefficients appear constant (Figure II.4). One step residuals remain within ±2 standard errors bands. One step Chow test also suggests constancy of estimated coefficients, although there is a spike in 2000. Recursive coefficients can be estimated backwards as well. The estimated coefficients appear somewhat uncertain around 2002-03, while as additional observations added backwardly, become more stable with confidence bands getting narrower. Similar results can be derived for model 7, as well (Figure II.5). Estimated coefficients seem broadly stable through the sample period, although in 2002-03, some parameter instability can be observed by backward recursive regression results. In sum, for both models, the parameter constancy tests confirm reasonably stable parameters, although there is some evidence of shift or uncertainty in 2002-03, which corresponds to a period when macroeconomic stability improved remarkably with the achievement of low inflation and stable exchange rates.

Estimated models 5 and 7 for Lao P.D.R. are:

$$\text{(model 5)}\ \ 0.059 - \underset{[-2.070]^{*}}{0.016} \times rlc_t - \underset{[-4.32]^{**}}{0.010} \times dcpi = 0$$

(model 7) $0.034 - \underset{[-1.200]}{0.009} \times rlc_t + \underset{[4.44]^{**}}{0.007} \times dneer = 0$

These estimated models can be simplified as:

(model 5) $rlc_t = 3.675 - 0.643 \times dcpi$

(model 7) $rlc_t = 3.847 + 0.795 \times dneer$

Recall that *dcpi* and *dneer* are calculated as monthly changes in CPI and nominal effective exchange rate (not log difference). The estimated coefficients in models 5 and 7 imply decreasing semi-elasticity of real monetary aggregates with respect to inflation and depreciation of exchange rates. To obtain the elasticity of real money balances with respect to those variables, we need some manipulations. Taking the antilogarithm of the equation (2) and ignoring *rgdp* terms, we get:

$$\frac{M}{P} = \exp(c + \beta_2 \times oc)$$

$$(\because) \quad \frac{\partial(M/P)}{\partial oc} = -\beta_2 \frac{M}{P}.$$

Hence, the elasticity in the absolute term is:

$$\left| \frac{\partial(M/P)}{M/P} \frac{oc}{\partial oc} \right| = \left| -\beta_2 \times oc \right|.$$

This implies that the elasticity increases as the *oc* increases. The elasticity of real currency with respect to change in the CPI and exchange reached nearly 16 in the late 1990s (Figure II.6) when Lao P.D.R. faced severe macroeconomic instability. This implies that one percentage point increase in inflation or exchange rate depreciation would have resulted in

a 16 percent decline in real local money stock. In recent years, the elasticities have fallen to a very low level, suggesting stability of real local money.

VI. Some Extensions To the Empirical Model For Lao P.D.R.

Because the parameter stability tests suggest some instability of estimated coefficients, I split the sample period to examine the robustness of the above results for Lao P.D.R. As discussed, Lao P.D.R. experienced quite serious macroeconomic instability during the late 1990s, but have achieved broad stability around 2001-02. Thus, I split the sample at the end-2000 and re-estimate ARDL models, followed by the ARDL bounds test.

ARDL bounds test suggest a quite contrast between these two sample periods (Table II.6). For the first half sample period, 1996 M2-2000 M12, there was clear evidence of the existence of long-run relationship between *rlc* and *dcpi* or *dneer* because the Wald test statistics were way above the upper critical value. However, for the second half sample period, the Wald test statistics fall below the lower critical value, not rejecting the null hypothesis of no long-run relationship.

The estimated models for 1996 M2-2000 M12 are:

$$\text{(model 5)}\quad 0.106 - \underset{[-2.20]^{*}}{0.038}\times rlc_t - \underset{[-3.42]^{**}}{0.013}\times dcpi = 0$$

$$\text{(model 7)}\quad 0.047 - \underset{[-1.11]}{0.018}\times rlc_t + \underset{[3.02]^{**}}{0.007}\times dneer = 0$$

By arranging these models, we can get the semi-elasticities of real local currency money at 0.350 for model 5 and 0.406 for model 7, smaller than those calculated for the full sample period (0.643 and 0.795, respectively). These are consistent with the previous results of the parameter stability tests, and could possibly reflect a structural shift between up to early 2000s (macroeconomic instability periods) and thereafter (relatively stable periods).

To further investigate a possible structural shift, Ie split the sample periods by introducing a dummy variable:

$$(5)\ \Delta rlc_t = c + \beta_1 \times D + \beta_1 \times rlc_{t-1} + \beta_2 \times oc_{t-1} + \beta_3 \times D \times rlc_{t-1} + \beta_4 \times D \times oc_{t-1}$$

$$+ \sum_{i=1}^{p} \gamma_i \Delta rlc_{t-i} + \sum_{k=0}^{r} \gamma_k \Delta op_{t-k} + \varepsilon_t$$

where c is the intercept term and ε_t are white noise errors. The equation (5) is equivalent to equation (3) except that $rgdp_{t-1}$ and $\sum_{j=0}^{q} \gamma_j \Delta rgdp_{t-j}$ have been omitted because my focus is the long run relationship between real money and opportunity costs. The structural change is captured by the dummy variable D:

$$D = \begin{cases} 1 & if\ t \le December\ 2000 \\ 0 & if\ t > December\ 2000 \end{cases}$$

The results suggest against the existence of the structural break (Table II.7) at the end of 2000. In model 5, while intercept dummy (β_1) and slope dummy for *drlc(-1)* (β_3) can be accepted at the five percent level of significance, the joint null hypothesis of $\beta_1 = \beta_3 = \beta_4 = 0$ and $\beta_3 = \beta_4 = 0$ cannot be rejected. The similar results hold for model 7, and there is no statistical evidence of structural break using the dummy variable.

VII. Policy Implications

One of the important policy implications that we can derive from the empirical results is that in Lao P.D.R., dollarization would be reversible provided that the authorities pursue sound and prudent macroeconomic policy. The evidence in models 5 and 7 suggest that if the macroeconomic policy makers are successful in gaining credibility to lower inflation and improve confidence in the exchange rate (i.e., appreciation), there will be an increased demand for kip, the Lao P.D.R's currency.

Unfortunately the recent performance of policy in the Lao P.D.R. raises concern. Since the middle of 2007, inflation pressures have remerged, and the headline inflation rates reached beyond 10 percent by the middle of 2008 (up from below 4 percent in the middle of 2007). This is due in part by global oil and food price trends, but is also attributable to lax control over aggregate demand. Monetary policy has been expansionary over the past year with credit growth reaching nearly 50% (year on year) by the middle of 2008. My results suggest that such an expansionary monetary policy—and as its consequence, higher inflation—would discourage the use of local currency money. This is inconsistent with the government's long-term desire to de-dollarize Lao P.D.R.'s economy.

On the contrary, for Cambodia, unfortunately, the same policy recommendation cannot be made based on the empirical results. There does not appear to be a long-run relationship between real local currency money and changes in inflation or changes in exchange rates. This may have been due to other factors, including extreme historical experiences.

VIII. CONCLUSIONS

My paper has presented stylized facts of the demand for local currency money in Cambodia and Lao P.D.R. Although both countries are located in geographically the same region, have shared the similar economic history, and are in the similar development stage, I have found a sharp contrast about statistical properties of the money demand between these two countries. I employed ARDL bounds test, rather than Johansen's cointegration approach, to find long-run demand for real local money, because the order of integration for the variables of interest is not clear,. The main findings of my research are paper are:

First, real local currency money and expected inflation or exchange changes in exchange rates have a long-run relationship in Lao P.D.R. However, real GDP is not a significant variable entered in the long-run relationship. Hence, a Cagan-type demand for money specification applies to Lao P.D.R. On the contrary, for Cambodia, such a relationship cannot be confirmed: the demand for local currency money does not demonstrate a long-run relationship with inflation or a change in the exchange rate. Although it is beyond the scope of my paper to prove this, we think that lack of use of local currency money or extensive degree of dollarization in Cambodia seems to reflect its unique historical factors: massive inflows of dollar cash during the early 1990s when the United Nation took over the country's administration, and maybe, people's yet weak trust to the country as their memory of the Khmer Rouge era remains.

Second, the long-run relationship between real local currency money and inflation or a change in exchange rate in Lao P.D.R. seems stronger in the late 1990s-early 2000s than thereafter periods. For the first half, there is statistically clear evidence of cointegration, but for the second half, we cannot confirm such evidence. However, other statistical tests, such as parameter stability tests and dummy variable approach, do not strongly support the existence of a structural break.

Finally, the results in my paper imply that people in Lao P.D.R. switch between holding local and foreign currency money depending on expected changes in inflation rates, but not in Cambodia. Thus, it might be possible for Lao P.D.R. to de-dollarize the economy, and promote the use of local currency money. This could be accomplished through effective monetary policy, if and only if they maintain macroeconomic stability. The policy implications do not directly apply to Cambodia, because there is limited supporting statistical evidence.

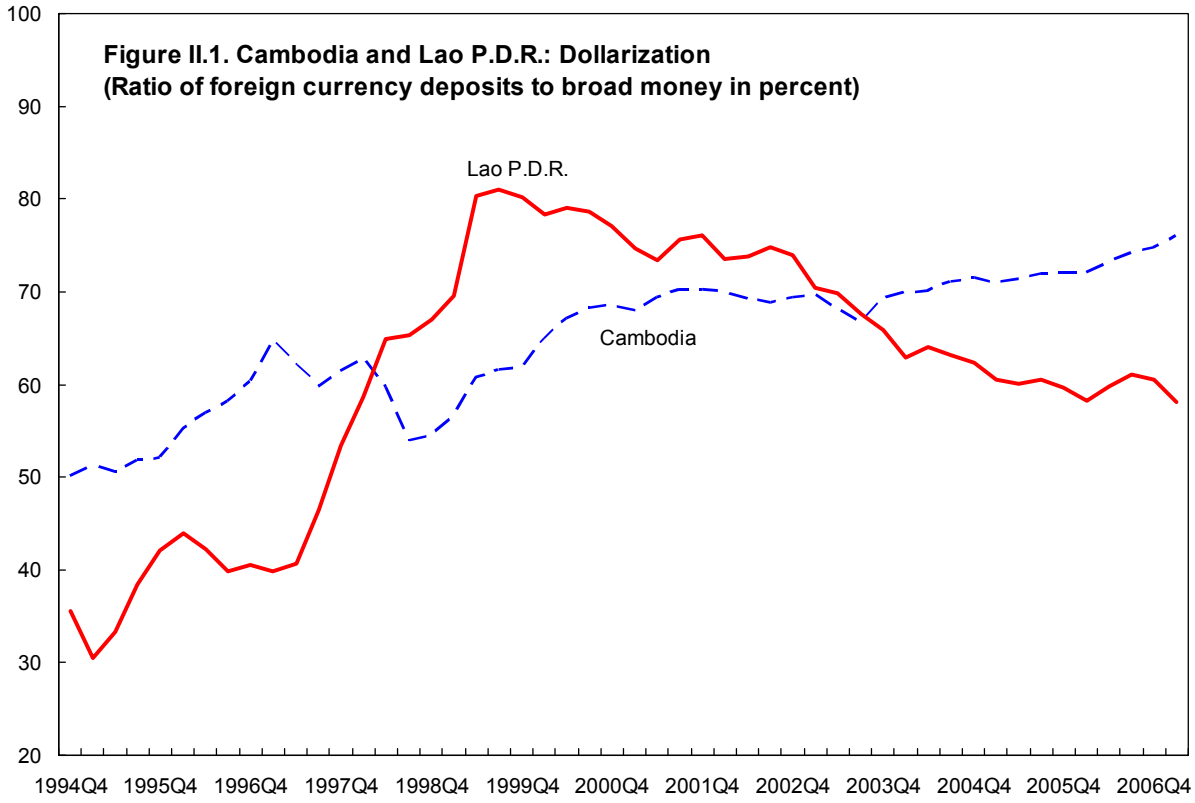


Figure II.1. Cambodia and Lao P.D.R.: Dollarization
(Ratio of foreign currency deposits to broad money in percent)
Lao P.D.R.
Cambodia
100
90
80
70
60
50
40
30
20
1994Q4
1995Q4
1996Q4
1997Q4
1998Q4
1999Q4
2000Q4
2001Q4
2002Q4
2003Q4
2004Q4
2005Q4
2006Q4

Figure II.2. Cambodia: Money, Price, and Exchange Rate

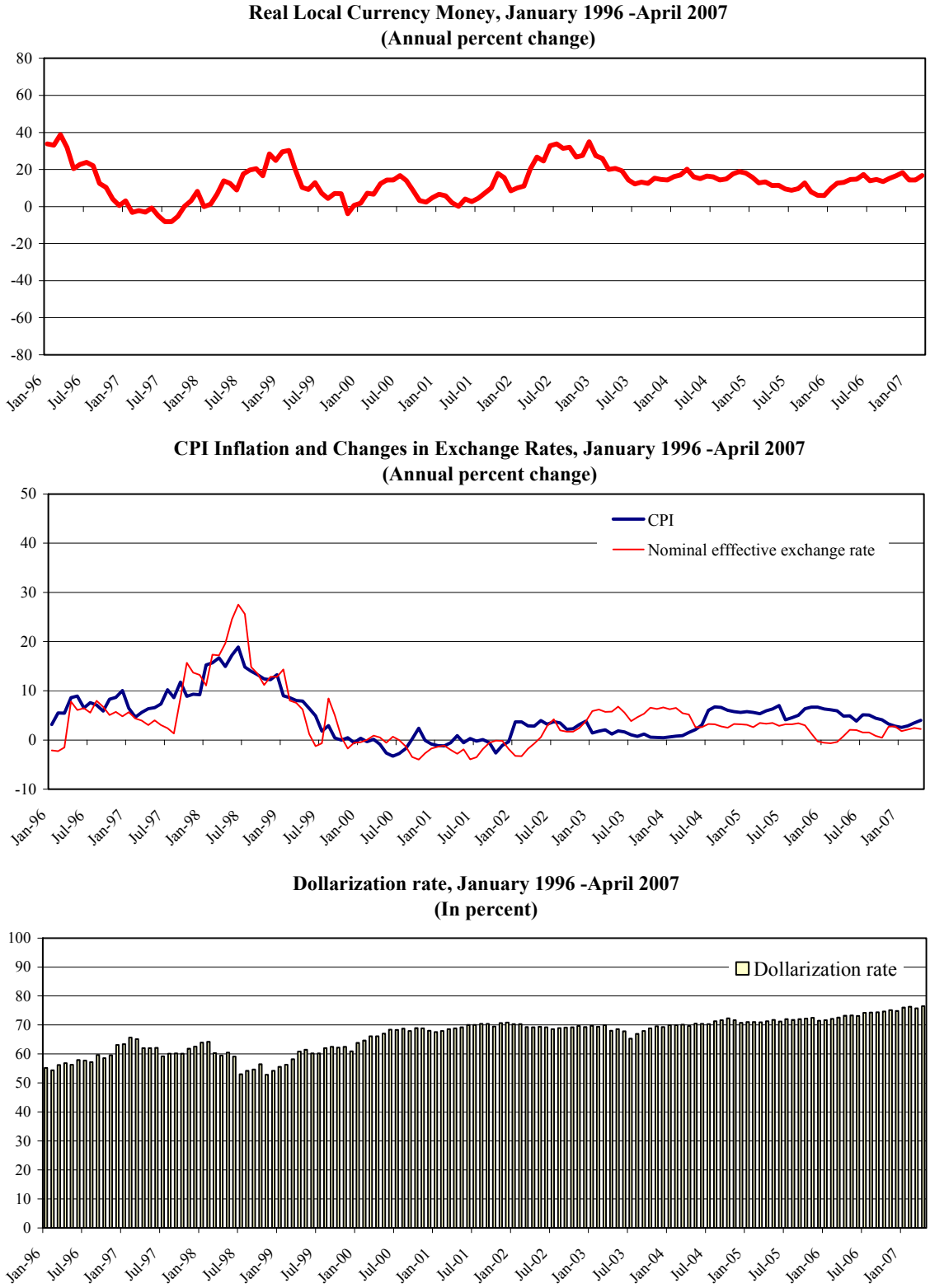


Figure II.3. Lao P.D.R.: Money, Price, and Exchange Rate

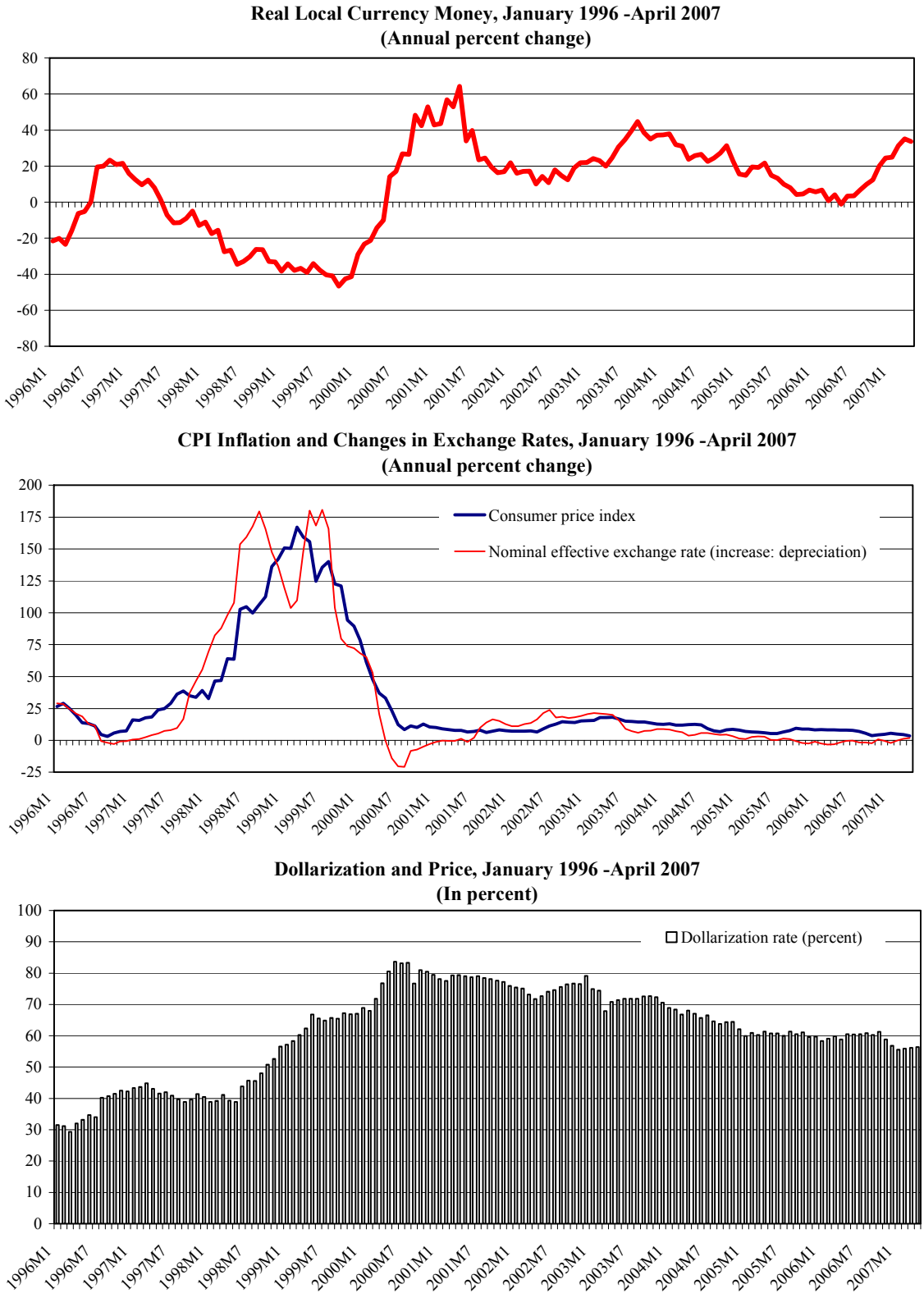


Figure II.4. Lao P.D.R. Model 5 (rlc and dcpi as a forcing variable) Recursive Estimation

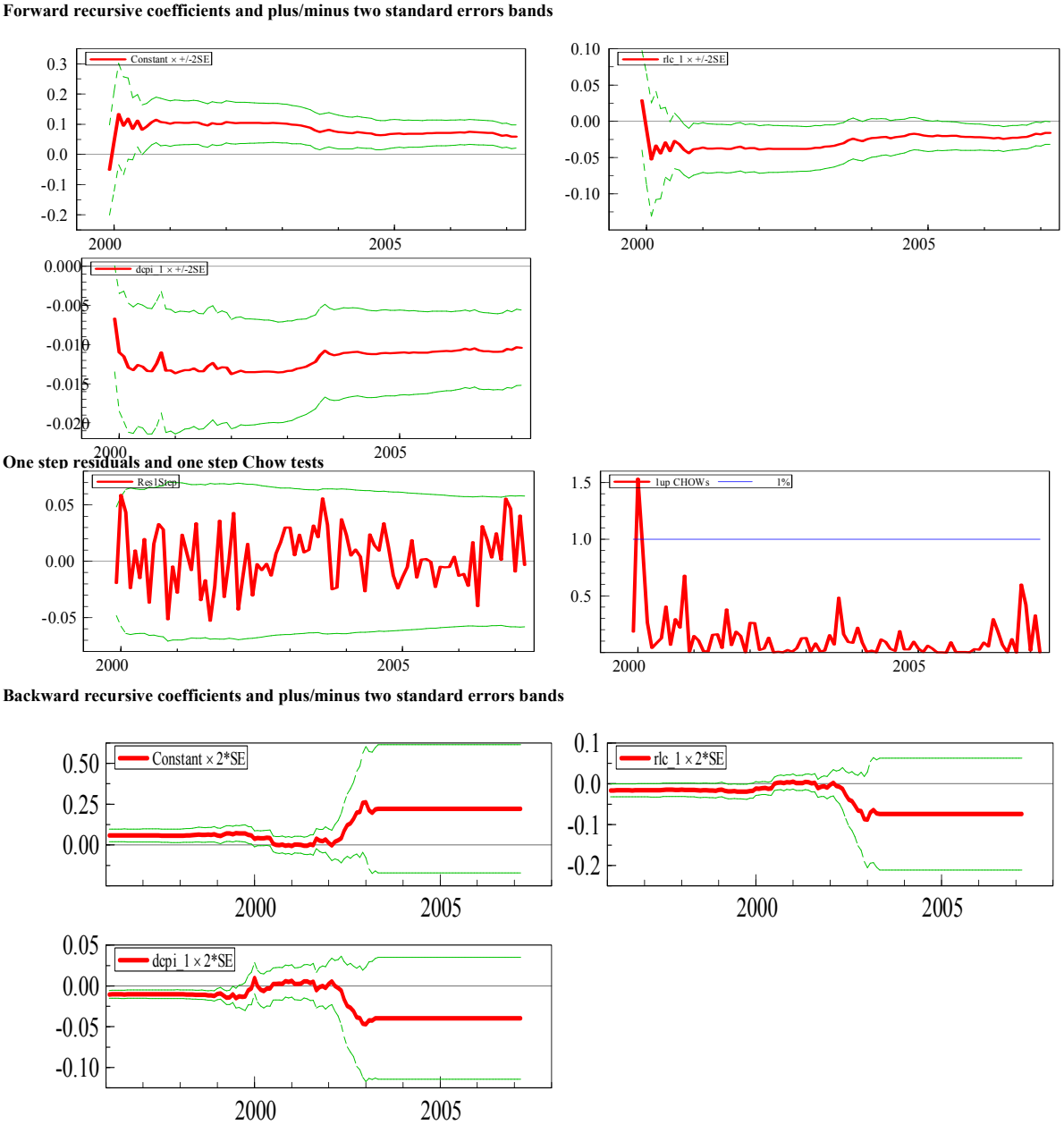
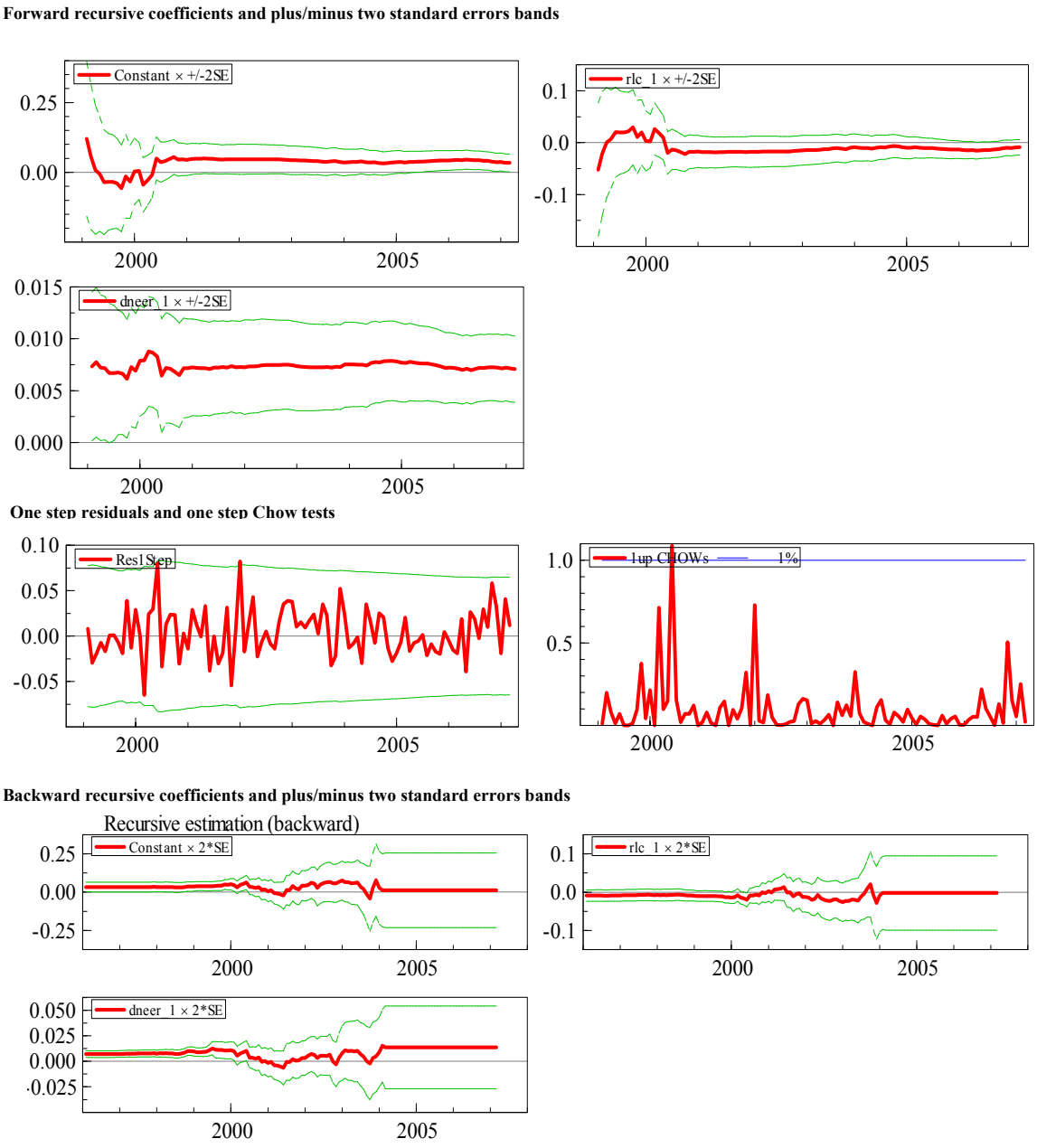


Figure II.5. Lao P.D.R. Model 7 (rlc and dneer as a forcing variable) Recursive Estimation



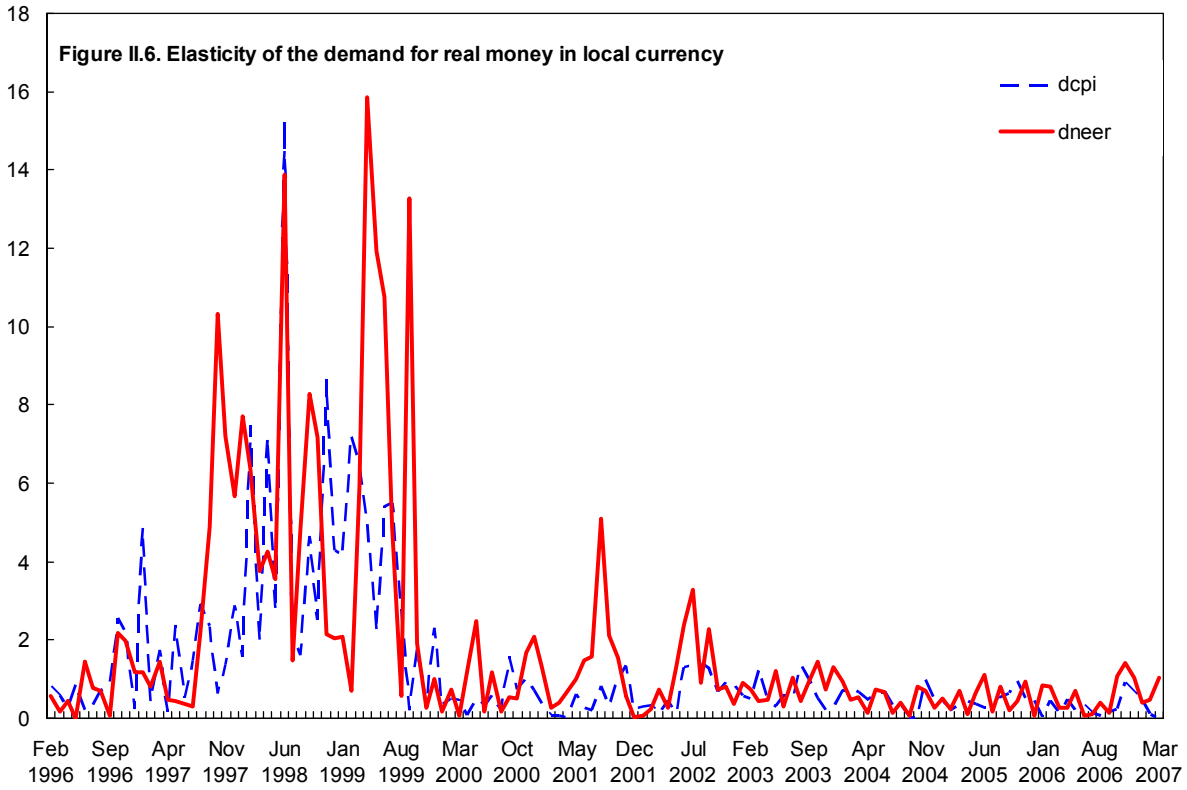


Figure II.6. Elasticity of the demand for real money in local currency
dcpi
dneer
18
16
14
12
10
8
6
4
2
0
Feb 1996
Sep 1996
Apr 1997
Nov 1997
Jun 1998
Jan 1999
Aug 1999
Mar 2000
Oct 2000
May 2001
Dec 2001
Jul 2002
Feb 2003
Sep 2003
Apr 2004
Nov 2004
Jun 2005
Jan 2006
Aug 2006
Mar 2007

Table II.1. Cambodia and Lao P.D.R.: Descriptive Statistics
(Sample periods: January 1995-March 2007)

	Mean	Standard deviation	Skewness
Cambodia			
Rlc (log)	1.908	0.437	0.236
RGDP (log)	9.612	0.295	0.242
Dcpi (monthly percent change)	0.349	0.900	0.974
Dneer (monthly percent change)	-1.578	3.995	-1.418
Lao P.D.R.			
Rlc (log)	2.076	0.400	-0.338
RGDP (log)	7.067	0.213	0.067
Dcpi (monthly percent change)	1.953	3.188	3.384
Dneer (monthly percent change)	-1.582	3.992	-1.413

Table II.2. Cambodia and Lao P.D.R. Unit root test
—Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) tests 1/
(Sample period: 1995M1 - 2007M3)

	Level				First difference				Comments on integration order
	Trend	Lag length *p*	ADF statistics 2/	Estimated γ	Trend	Lag length *p*	ADF statistics 2/	Estimated γ	
				Augmented Dickey-Fuller Test					
Cambodia									
Rlc (log)	No	0	-0.220	0.999	No	0	-12.584**	-0.047	I(1)
RGDP (log)	No	12	2.799	1.003	No	12	-2.708	0.754	I(2) possible
Dcpi (monthly percent change)	No	0	-12.263**	-0.015	No	4	-10.379**	-3.035	I(0)
Dneer (monthly percent change)	No	0	-6.162**	0.583	No	5	-9.647**	-1.917	I(0)
Lao P.D.R.									
Rlc (log)	No	5	-0.729	0.994	No	4	-2.873	0.655	I(2) possible
RGDP (log)	No	12	2.195	1.001	No	12	-3.897**	0.473	I(1)
Dcpi (monthly percent change)	No	3	-2.328	0.804	No	3	-10.242**	-2.083	I(1)
Dneer (monthly percent change)	No	0	-6.095**	0.590	No	5	-9.556**	-1.900	I(0)
				Phillips-Perron Test					
Cambodia									
Rlc (log)	No	...	-0.213	0.999	No	...	-12.577**	-0.047	I(1)
RGDP (log)	No	...	3.175	1.005	No	...	-7.276**	0.544	I(1)
Dcpi (monthly percent change)	No	...	-12.317**	-0.015	No	...	-69.900**	-0.572	I(0)
Dneer (monthly percent change)	No	...	-6.176**	0.583	No	...	-29.847**	-0.283	I(0)
Lao P.D.R.									
Rlc (log)	No	...	-0.248	1.006	No	...	-15.296**	0.213	I(1)
RGDP (log)	No	...	0.973	1.001	No	...	-9.123**	0.346	I(1)
Dcpi (monthly percent change)	No	...	-9.149**	0.429	No	...	-39.242**	-0.614	I(0)
Dneer (monthly percent change)	No	...	-6.104**	0.590	No	...	-29.481**	-0.260	I(0)

1/ For a variable y, the following regression equation is considered:

ADF test: $$\Delta y_t = a_1 + a_2 \cdot t + (\gamma - 1) \cdot y_{t-1} + \sum_{i=1}^{p} \beta_i \Delta y_{t-i} + \varepsilon_t$$

PP test: $$\Delta y_t = a_1 + a_2 \cdot t + (\gamma - 1) \cdot y_{t-1} + \nu_t$$

where Δ is the difference operator, t a time trend, and p the lag length. The ADF and PP statistics are the t ratio on the lagged variable .

2/ The null hypothesis is the variable has a unit root. ** indicates significant at 1 percent level and * indicates significant at 5 percent level. The null hypothesis is γ =1 (no unit root).

Table II.3. ARDL Model: Lag Selection

(Sample period: 1996M2-2007M3)

Forcing variable(s)	*Inflation*			*Inflation and real GDP*			*Exchange rate*			*Exchange rate and real GDP*		
	F statistics	AIC	SC	F statistics	AIC	SC	F statistics	AIC	SC	F statistics	AIC	SC
						(Cambodia)						
Lag 12	...	-4.599	-3.993	...	-4.638	-3.730	...	-4.684	-4.078	...	-4.762	-3.853
Lag 11	1.923	-4.593	-4.031	1.193	-4.645	-3.802	4.336	-4.635	-4.073	0.301	-4.797	-3.953
Lag 10	1.030	-4.604	-4.085	0.266	-4.681	-3.903	0.064	-4.664	-4.145	0.523	-4.825	-4.046
Lag 9	0.580	-4.623	-4.148	2.103	-4.664	-3.950	4.538	-4.614	-4.138	6.2464**	-4.695	-3.981
Lag 8	0.385	-4.646	-4.214	0.268	-4.701	-4.052	0.188	-4.641	-4.208	0.908	-4.713	-4.064
Lag 7	1.174	-4.656	-4.267	0.912	-4.719	-4.136	1.577	-4.643	-4.254	0.909	-4.732	-4.148
Lag 6	0.513	-4.677	-4.331	0.387	-4.755	-4.234	0.234	-4.669	-4.323	0.597	-4.760	-4.241
Lag 5	0.982	-4.690	-4.387	0.985	**-4.772**	-4.318	0.188	**-4.696**	-4.392	0.525	**-4.791**	-4.336
Lag 4	1.719	-4.692	-4.432	2.8697*	-4.743	-4.354	**8.226****	-4.597	-4.338	7.8401**	-4.646	-4.257
Lag 3	0.234	**-4.718**	-4.502	0.997	-4.762	-4.438	0.044	-4.626	-4.410	0.207	-4.686	-4.361
Lag 2	**4.622***	-4.676	-4.503	**4.9127****	-4.690	-4.431	1.405	-4.634	-4.461	**3.2031***	-4.653	-4.393
Lag 1	0.397	-4.699	**-4.570**	2.351	-4.679	**-4.484**	1.039	-4.647	**-4.517**	2.640	-4.635	**-4.440**
						(Lao P.D.R.)						
Lag 12	...	-4.078	-3.472	...	-4.039	-3.131	...	-3.831	3.225	...	-3.887	-2.979
Lag 11	2.614	-4.060	-3.498	**4.278****	-3.953	-3.110	1.341	-3.835	-3.273	0.773	-3.907	-3.063
Lag 10	0.306	-4.084	-3.565	0.200	-3.992	-3.213	0.141	-3.862	-3.344	0.129	**-3.947**	-3.169
Lag 9	1.263	**-4.091**	-3.615	0.778	-4.013	-3.299	0.677	-3.880	-3.405	1.606	-3.944	-3.231
Lag 8	1.815	-4.089	-3.657	1.762	-4.007	-3.358	3.272	-3.853	-3.421	2.926*	-3.906	-3.257
Lag 7	1.788	-4.088	-3.700	1.344	-4.014	-3.430	0.363	-3.878	-3.488	0.997	-3.922	-3.338
Lag 6	4.548*	-4.042	-3.696	1.424	-4.019	-3.500	0.556	-3.897	-3.551	2.422	-3.901	-3.382
Lag 5	0.097	-4.071	-3.768	0.614	**-4.047**	-3.593	0.191	**-3.924**	-3.621	0.420	-3.934	-3.480
Lag 4	**5.699****	-4.010	-3.750	4.840**	-3.971	-3.582	**6.239****	-3.855	3.595	**4.425****	-3.868	-3.479
Lag 3	0.308	-4.035	-3.818	1.060	-3.989	-3.665	5.942**	-3.792	-3.575	3.710	-3.821	-3.497
Lag 2	1.634	-4.038	-3.865	1.000	-4.009	-3.749	2.208	-3.786	**-3.613**	2.172	-3.813	-3.553
Lag 1	1.728	-4.041	**-3.989**	0.842	-4.033	**-3.839**	7.563**	-3.703	-3.573	4.021**	-3.763	**-3.569**

Table II.4. ARDL model diagnostic tests 1/
(Sample period: 1996M2-2007M4)

Forcing variables	Cambodia				Lao P.D.R.			
	Inflation	*Inflation and real GDP*	*Exchange rate*	*Exchange rate and real GDP*	*Inflation*	*Inflation and real GDP*	*Exchange rate*	*Exchange rate and real GDP*
Model number	*1*	*2*	*3*	*4*	*5*	*6*	*7*	*8*
Selected lags	**3**	**5**	**5**	**5**	**9**	**12**	**5**	**10**
Autocorrelation LM test 2/	0.925	0.398	0.121	0.580	1.714	1.532	0.978	0.347
Probability	*0.49*	*0.90*	*1.00*	*0.77*	*0.11*	*0.17*	*0.45*	*0.93*
Normality 3/	28.97	44.58	1.275	4.55	3.56	1.46	5.804	0.310
Probability	*0.00* **	*0.00* **	*0.53*	*0.10* **	*0.17*	*0.48*	*0.05*	*0.86*
Heteroscedasticity 4/	1.03	0.79	1.856	1.80	1.92	0.26	0.976	0.778
Probability	*0.43*	*0.78*	*0.02* *	*0.02* *	*0.01* **	*1.00*	*0.51*	*0.80*

Notes:
1/ ** indicates significant at 1 percent level and * indicates significant at 5 percent level.
2/ The Lagrange-multiplier test for 7th order residual autocorrelation, distributed as Chi squares (7) in large samples. Under the null hypothesis that there is no autocorrelation.
3/ Normality is Doornik and Hansen normality (1994) test. Under the null hypothesis, residuals are normal distributed.
4/ Heteroscedasticity is Doornik (1996) test. Under the null hypothesis, there is no heteroskedasticity.

Table II.5. ARDL Bound Tests for the Existance of a Long-run Relationship and Coefficients 1/ 2/
(Sample period: 1996M2 - 2007 M3)

	Cambodia				Lao P.D.R.			
Model number	Inflation 1	Inflation and real GDP 2	Exchange rate 3	Exchange rate and real GDP 4	Inflation 5	Inflation and real GDP 6	Exchange rate 7	Exchange rate and real GDP 8
ARDL bound tests 3/								
Wald statistics	2.197	11.875	1.215	15.140	18.722	9.812	20.303	14.450
Evidence for long-run relatonship?	No	Undetermined	No	Yes	Yes	No	Yes	Undetermined
Long-run coefficients								
Constant	0.003	-1.006	-0.004	-0.945	0.059	-0.043	0.034	-0.091
	0.302	*-2.490 **	*-0.349*	*-2.490 **	*3.09 ***	*-0.337*	*2.14 **	*-0.755*
rlc (-1)	0.004	-0.069	0.006	-0.057	-0.016	-0.010	-0.009	-0.007
	0.868	*-2.030 **	*1.100*	*-1.780*	*-2.070 **	*-0.798*	*-1.200*	*-0.802*
dcpi (-1)	-0.004	0.005	...	...	-0.010	-0.011	...	...
	-0.993	*0.874*			*-4.32 ***	*-3.080 ***		
dneer (-1)	...	...	-0.001	-0.001	...	...	0.007	0.010
			-0.656	*-1.590*			*4.44 ***	*3.530 ***
rgdp (-1)	...	0.120	...	0.112	...	0.015	...	0.026
		*2.490 **		*2.470 **		*0.829*		*1.470*
Memorundam item:								
Number of lags	2	5	5	5	9	12	5	10

Notes:
1/ *Italic* indicates t statistics; [ ] indicates p value.
2/ ** indicates significant at 1 percent level and * significant at 5 percent level.
3/ Critical value used for bounds Wald statistics at a five percent significance level
With constant and one forcing variables: upper bound I(1), 11.528; lower bound I(0), 9.867.
With constant and two forcing variables: upper bound I(1), 14.566; lower bound I(0), 11.380.

Table II.6. Lao P.D.R.: ARDL Bound Tests for Sub-sample Periods 1/ 2/

Model number	Model 5 (inflation as a forcing variable)			Model 7 (exchange rate as a forcing variable)		
	1996M2-2007M3 (full sample)	1996M2-2000M12	2001M1-2007M3	1996M2-2007M3 (full sample)	1996M2-2000M12	2001M1-2007M3
ARDL bound tests 3/						
Wald statistics	18.722	12.525	0.552	20.303	12.951	0.552
Evidence for long-run relationship?	Yes	Yes	No	Yes	Yes	No
Long-run coefficients						
Constant	0.059	0.106	-0.007	0.034	0.047	0.005
	3.09 **	*2.80* **	*-0.26*	*2.14* *	*1.61*	*0.16*
rlc (-1)	-0.016	-0.038	0.006	-0.009	-0.018	0.002
	-2.07 *	*-2.20* *	*0.59*	*-1.200*	*-1.11*	*0.18*
dcpi (-1)	-0.010	-0.013	0.006	...	...	...
	-4.32 **	*-3.42* **	*0.63*			
dneer (-1)	...	...	...	0.007	0.007	-0.002
				4.44 **	*3.02* **	*-0.26*
Diagnostic tests						
				...		
Autocorrelation LM test	1.714	1.469	1.302	0.978	1.868	0.951
Probability	*0.11*	*0.23*	*0.28*	*0.45*	*0.13*	*0.46*
Normality	3.562	0.809	0.209	5.804	2.945	0.852
Probability	*0.17*	*0.20*	*1.90*	0.05	*0.23*	*0.65*
Heteroscedasticity	1.922	49.559	0.874	0.976	0.274	0.903
Probability	*0.01* **	*0.20*	*0.65*	*0.51*	*1.00*	*0.60*
Memorandum item:						
Number of lags	9	9	9	5	9	9
Number of observations	134	59	75	134	59	75

Notes:
1/ *Italic* indicates t statistics; [ ] indicates p value.
2/ ** indicates significant at 1 percent level and * significant at 5 percent level.
3/ Critical value used for bounds Wald statistics at a five percent significance level
With constant and one forcing variables: upper bound I(1), 11.528; lower bound I(0), 9.867.
With constant and two forcing variables: upper bound I(1), 14.566; lower bound I(0), 11.380.

Table II.7. Lao P.D.R.: Dummy Variable Approach 1/ 2/

Model number	Model 5 (inflation as a focing variable)	Model 7 (exchange rate as a forcing variable)
Long-run coefficients		
Constant	0.038 *1.59*	0.039 *1.68*
D	0.057 *2.11* *	0.010 *0.38*
rlc (-1)	-0.008 *-0.86*	-0.010 *-1.01*
D×rlc (-1)	-0.027 *-2.16* *	-0.010 *-0.80*
dcpi (-1)	-0.007 *-1.11*	...
D×dcpi (-1)	-0.004 *-0.80*	...
dneer (-1)	...	0.007 *2.16* *
D×dneer (-1)	...	-0.001 *-0.28*
Joint null hypotheses		
D, D×rlc (-1), and D×dneer (-1)=0	5.527 [0.137]	2.421 [0.490]
D×rlc (-1) and D×dneer (-1)=0	5.440 [0.066]	0.730 [0.694]
Memorundam items:		
Number of lags	9	5
Number of observations	134	134

Notes:
1/ *Italic* indicates t statistics; [ ] indicates p value.
2/ ** indicates significant at 1 percent level and * significant at 5 percent level.

Essay III: Modeling Electricity Demand in the Philippines

I. Introduction

Because power is one of the most essential goods for production and consumption, an adequate, efficient, and reliable supply is critical for sustainable growth and welfare. In the developing country growth literature, there is a broad consensus that electricity generation and GDP growth appear closely correlated (Calderón and Servén, 2004, and IEA, 2004). For example, the Philippines faced chronic power shortages from the late 1980s to early 1990s. The most severe power crisis occurred in 1992-93. The poor economic performance has been attributed in part to the power crises.

Governments in developing countries are often responsible for planning electricity sector development, including the demand forecasts. There are significant costs and risks from making forecast errors. However, this poses a major challenge to policymakers.

First, under forecasting electricity consumption will lead to insufficient investment in the power sector, leading to risks of blackouts and is a drag on economic activity. Second, given the long lead time and lumpiness for constructing future generating capacity, the electricity shortages will persist after becoming apparent. If on the other hand, government planners over-predict electricity consumption, the costs come in two forms. There are significant financing and borrowing costs associated with infrastructure expenditures like power generation and networks. The debt servicing costs are long-term. Developing countries are already constrained financially. Large long-term financial obligations with no

return represent lost opportunities and a waste of limited public resources. Despite the importance of electricity demand forecasting, there have been relatively few empirical studies on electricity demand for developing countries, including the Philippines.

The objective of my research is to model electricity demand in the Philippines. The analysis is based on examining the long-run and short-run determinants of residential and industrial consumption. The estimated price elasticity, income elasticities, economic development responses can provide useful information for the Philippine power sector development plan. The remainder of the paper is organized as follows: Section II summarizes the Philippines' economic and power sector developments over the last few decades. Section III provides a brief literature review on electricity and econometric modeling. Section IV describes the modeling approach and data. Section V presents the long-run electricity demand model results, and Section VI incorporates these into an electricity demand model with short-run dynamics. Section VII discusses implications of the estimated models in the context of power sector policy in the Philippines, and Section VIII concludes.

II. The Philippine Economy and Power Sector

Until the mid-1980s, the power sector in the Philippines was mostly state-owned through the National Power Corporation (NPC). After the debt crisis in the early 1980s, the Philippines' government tightened fiscal policy, and capital expenditure for additional electricity capacity was significantly reduced. This led to a slowdown in the electricity generating facilities. At the same time, electricity demand continued to increase. This

resulted in tight electricity supply and demand conditions by the middle of the 1980s (Figure III.1).

The government decided to encourage the private sector to enter the electricity market. To reduce risks of emerging electricity shortages, President Corazón Aquino issued Executive Order 215 of 1987, breaking the NPC's monopoly and allowing private investors to enter into the generation market. However, private participation did not increase as much as intended. Electricity shortages led to a power crisis in 1992-93. During the crisis period, power outages were several hours a day in many areas of the country (De Dios and Hutchcroft, 2003). To resolve the crisis, the government approved the Electricity Power Crises Act of 1993, granting the President special powers to solve the energy crisis, including those to approve power projects. With such powers in hand, President Fidel Ramos further deregulated generation markets to encourage private investors and actively courted independent power producers (IPPs).

In 1997, the Philippines was hit by a major external shock, the Asian financial crisis. Electricity demand declined sharply from 10 percent growth per year in 1994-97 to only two percent in 1998-99. There was suddenly a large excess supply in the electricity market. Under the power purchase agreements with IPPs, NPC had to continue purchasing the contracted amount of power from IPPs, which were not produced or consumed. As a result, the NPC started running a loss every year from 1998, which reached over 1 percent of GDP by 2003.

Electricity demand has picked up since 2000 with the economic recovery. Today, the total installed generation capacity amounts to 15,620 MW, of which 20 percent are from hydro plants, 50 percent from oil, diesel, and coal plants, and 30 percent from geothermal and natural gas plants. The Philippine Department of Energy predicts that power generating capacity will need to increase by 40 percent (5,200 MW) by the 2014 to avoid the power shortages. In 2001, the Philippines' government launched a comprehensive power sector reform plan, including market liberalization and NPC privatization.[1] Again, the goal was to increase private investment and participation in the electric power sector.

III. Literature Review

There have been numerous of studies on electricity demand for advanced countries. These studies typically attempted to estimate long- and short-run elasticities of electricity consumption with respect to income and price. They are summarized in Table III.1. Atkinson and Manning (1995) surveyed the elasticities estimated in the literature before the early 1990s. The studies of recent years include those for the U.S. (Silk and Joutz, 1997; Kamerschen and Porter, 2004), U.K. (Clements and Madlener, 1999), Australia (Akmal and Stern, 2001; Narayan and Smyth, 2005), New Zealand (Fatai et al., 2003), and Switzerland (Filipini, 1999).

On the contrary, econometric studies on electricity demand, addressing both long- and short-run elasticities, in developing countries, have been relatively limited. Pesaran et al

[1] This reform initiative is embodied in the Electricity Power Industry Reform Act (EPIRA) enacted in 2001.

(1998) investigated electricity demand in ten Asian economies, Holtedahl and Joutz (2004) for Taiwan, Erdogan and Dahl in Turkey (1997), Lin in China (2003), Galindo in Mexico (2005), Zachariadis et al (2006) in Cyprus, De Vita et al (2005) in Namibia. De Vita et al. (2005) surveyed studies on electricity demand in developing economies. Research on electricity consumption in developing economies is hindered by the difficulty in obtaining consistent data for time series analysis.

However, there are relatively large number of studies investigating causal relationships between electricity and growth. Foe example, Cheng and Lai (1997) and Yang (2000) investigated the direction of causality between energy consumption and growth in Taiwan, Glasure and Lee (1997) in South Korea and Singapore, Masih and Masih (1997) in Korea and Taiwan, Morimoto and Hope (2004) in Sri Lanka, Yoo (2005) in Korea, Asafu-Adjaye (2000) in India, Indonesia, Thailand, and the Philippines, Ghosh (2002) in India, Shiu and Lam (2004) in China, and Jumbe (2004) in Malawi. Wolde-Rufael (2006) analyzed the causality using data for 17 African countries. However, these studies do not develop a fully-specified demand model, which includes long- and short-run dynamics of electricity demand.

To my knowledge, the only formal econometric analysis of electricity demand focusing on the Philippines which includes the pre- and post-Asian crisis periods was conducted by the Philippines' Department of Energy (DOE) in 2004 and 2005. Their analysis was based on the combination of three different methodologies by different entities involved in the DOE's power development planning. First, the DOE requires each distribution utility to

submit a distribution development plan, in which a distribution utility presents its medium-term demand forecast. Each has different forecasting methodologies. Typically, the electricity consumption growth drivers are based on projected number of customers by sector (residential, industrial, and commercial) and per-consumer electricity consumption. Second, a nation wide electricity transmission company, National Transmission Corporation (Transco), submits to the DOE its own demand forecast by grid—in the Philippines, the power grid consists of three different grids, namely Luzon grid (north island), Visaya grid (a group of central islands), and Mindanao grid (south island). Transco uses an engineering and econometric approach for demand forecast, and for the latter econometric approach, real GDP and population are main explanatory variables. Third, the NPC—the state-owned power generation company with the largest market share in the nation—also produces demand forecast based on a trend analysis. Finally, the DOE has the responsibility to put together all these inputs from distribution companies, the Transco, and the NPC, and after consultations with these entities, finalize the demand forecast and a power sector development plan.

My research attempts to contribute to improve the econometric approach conducted by Transco. Transco's econometric approach apparently neither addresses issues related to non-stationarity and integration nor takes account of long-run economic demand and its determinants. Accordingly, my contribution is to demonstrate how the cointegration methodology would be useful to model the Philippines' electricity demand.

IV. Model and Data

A. Model

Demand for electricity is estimated separately for residential and industrial (including mining, manufacturing, construction, utility, and services) sectors. The agriculture sector is excluded from the analysis because its electricity consumption is very small. Following Silk and Joutz (1997) and Holtedahl and Joutz (2004), long-run residential electricity demand can be presented as:

$$E_t^H = H(I_t, P_t^H, A_t)$$

where E_t^H is residential electricity consumption per capita at time t, I_t is real personal disposable income per capita, and P_t^H is the relative residential electricity price with respect to GDP deflator with a base year of 1985.

A_t is the real stock of electronic appliances per capita, which in the long-run, could be affected by the economic factors, as consumers adjust the level of the stock of appliances dependent on their income and prices. The collection and measurement of energy using capital stock is a typical problem in nearly all electricity consumption models. The proxy variable considered in my research is household expenditures on furnishings per capita. Although this measure includes non-appliance goods that do not require electricity for use, due to limited data availability, this can be considered as the best available proxy for the stock of electronic appliances. My conjecture is that furnishings and electrical appliances

move together. The stock data are estimated by applying a perpetual inventory method to the household furnishings data (see below for more details).

Electricity consumption is expected to increase with income and the stock of electronic appliances and to decline if the price goes up. In a log linear form, the model that we estimate can be expressed as:

$$eh = c + \beta_1 \times i + \beta_2 \times ph + \beta_3 \times a \; ; \; \beta_1 > 0, \beta_2 < 0, \beta_3 > 0. \qquad (1)$$

where *c* is a constant term.

The industrial consumption model is derived from a production function. Electricity is an important input for the industrial sector. Let a production function be:

$$Y_t^P = F(K_t^P, E_t^P)$$

where Y_t^P is industrial output (GDP in industry and the services sector), K_t^P capital, and E_t^p is industrial electricity consumption. The variables are in per labor terms. Profit maximization behavior leads to the demand for the input electricity. This can be expressed in general terms as:

$$E_t^P = f(Y_t^P, P_t^P)$$

where P_t^P is the relative industrial electricity price with respect to the GDP deflator. Electricity demand is expected to increase as industrial output increases and price declines. In a log linear form, the model to be estimated is:

$$ep = c + \beta_1 \times gdp + \beta_2 \times pp\,;\ \beta_1 > 0, \beta_2 < 0. \tag{2}$$

B. Data

There are two data sources for electricity consumption (E^H and E^P) in the Philippines: data for the nationwide Philippines compiled by DOE and those for Manila areas compiled by Meralco (a distribution company). All data are annual. There are pros and cons from using either of these data sets. DOE data cover the whole nation, but sectoral data are available only from 1986 (aggregate data are available from 1971), limiting the scope for time series analysis. On the contrary, Meralco data are available from 1946 for sectoral consumption, but the data cover only Manila areas. Macroeconomic data that are determinants of electricity demand, such as income and GDP, for Manila areas are available only from 1986, not enough time span for modeling. Weighing on these pros and cons, I decide to use DOE data, but sectoral electricity consumption for the 1971-85 period is estimated using annual growth of the Meralco data.[2]

[2] I am grateful to then Undersecretary Ocampo and Ms. Exempe (both Philippine Department of Energy) for providing data to me and for their helpful discussions about this topic.

Data for population and labor, to calculate residential electricity consumption per capita and industrial electricity consumption per labor, are available from the Philippine National Statistical Coordination Board.

Electricity prices (P^H and P^P) are Meralco's sale prices. As discussed earlier, Meralco's franchise covers only the Metro Manila area. These prices may not be accurate representations for electricity prices throughout the rest of the Philippines. However, Meralco has maintained a dominant market share in the Philippines' electricity market (recently about 50 percent, and much higher in the past), and thus, the data can be considered as the best available proxy for the nationwide electricity price.

Data for real personal disposable income (*I*) and output (*GDP*) are from the Philippine National Accounts. Personal disposable income is available only from 1981, and data for pre-1981 periods need to be estimated. Because personal disposable income and private consumption are highly correlated, the growth rate of the private consumption is used to estimate the personal disposable income for the 1971-80 periods.

Data for household furnishings (*A*), a proxy variable for electric appliances, are from the Philippine National Accounts, but are only available since 1981. Because much longer time series are needed to estimate the stock of household furnishings, the growth rate of private consumption is used to estimate data for the household furnishings for the 1967-80 periods. Then, using a perpetual inventory model, the stock data are estimated. For this calculation,

a geometric depreciation rate of 0.3 is assumed, implying that the value of appliance would become nearly zero in ten years. Nominal values for household furnishings are deflated by the U.S. producer price index for household appliances[3]

Weather conditions may affect electricity use. The Philippines is a tropical country with relatively moderate weather conditions, but often face very high temperatures. Cooling degree-days are defined as the measurement of the severity and duration of hot weather and calculated as the difference between an average daily temperature and a reference temperature. If the former is lower or equal to the latter, a zero value is assigned. I set the reference level for the temperature at, 25ºC. This was selected because temperatures in major shopping malls in the Philippines have been set around 25ºC in recent years in compliance with the government energy saving policy. See Figure III.2.

Figures III.3.a, 3.b, and 3.c present the log levels of residential electricity consumption per capita with real disposable income per capita, real residential prices, and stock of household furnishings per capita respectively. The residential electricity consumption grew in tandem with real disposable income and electronic appliances until the early 1990s. Since then, electricity consumption has grown faster than income, but at a slower pace than appliances. The two factors may have offsetting effects. Residential electricity prices have

[3] In the Philippines, many of the items in household furnishings, such as electric appliances, are imported, in the early periods mainly from the U.S. and latter periods increasingly from the neighboring Asian countries. I use the U.S. price as proxy for the price of household furnishings. The alternative choice would be GDP deflator. However, I did not use it, because due to rapid technological growth in producing household furnishing goods, the price of household furnishing goods has grown at a slower pace than the overall GDP deflator in many countries.

steadily risen. The bivariate relationship seems inconsistent with the pattern of electricity consumption growth.

Figures III.4 shows graphs of the log levels for industrial electricity consumption per labor, output per labor, and prices. The industrial electricity consumption and output appear closely correlated until around 1987. During the 1990s, electricity consumption was more volatile (partly reflecting power crises) than output. Since 2000, it appears that growth of output is higher than that of electricity consumption. The price has been generally more volatile than electricity consumption, and the extent of correlation between the two variables is not immediately clear.

V. Econometric Modeling Strategy and Estimation of Long-run Electricity Demand

Econometric modeling strategy

I interpret a long-run electricity demand model, like equation (1) and (2), as describing an equilibrium relationship among the variables of interest (e.g., *eh*, *i*, *ph*, and *a* or *ep*, *gdp*, and *pp*). The empirical the model suggests a steady state relationship among these variables, given that these variables are growing over time. In other words, if these variables are linked to form an equilibrium relationship in the long-run, they will move closely together over time.

Such an interpretation is consistent with the objectives of my research because my main interest lies in the behavior of electricity consumption and its interactions with other

relevant variables (e.g., *i* or *gdp*) in the long-run. In this context, my research attempts to test whether an equilibrium relationship exists; and if so, whether it is stable.

Accordingly, the cointegration approach, suggested by Engel and Granger (1987) and Johansen (1988 and 1991), would be most suited for my research objectives. Between Engel and Granger's single equation approach and Johansen's multivariate system approach, I choose the latter approach. This is primarily because Johansen's VAR approach allows me to build a system or model without imposing any assumptions about endogeneity issues: whether changes in any of the right-hand-side variables in equation (1) or (2) cause changes in the left-hand-side variable *eh* or *ep*. In addition, I will employ a general-to-specific approach: I start with an explicit stochastic formulation of all *relevant* data in a system and then reduce the general dynamic model by imposing testable restrictions on the parameters. Here, the relevant data are chosen based on my theoretical interest and hypothesis—e.g., whether the electricity price would be associated with long-run electricity consumption— and then I will answer these questions based on a statistical analysis of the VAR model. Finally, I am also interested in the short-run dynamics of the variables under consideration. Especially since the steady state may rarely be observed—for example, the adjustment back to steady state is disturbed by new shocks around that level. In sum, I will attempt to develop an empirical model where the long-run determinants of electricity consumption by sector are accounted for and the short run determinants by an error correction model and or Autoregressive Distributed Lag (ARDL) model. The application of these techniques to electricity demand modelling has been fully

developed by Cheung et al (2003), Holtedahl and Joutz (2004), Halicioglu (2007), and Zachariadas et al (2007).

Estimation of long-run electricity demand

The first step is to analyze the times series properties of the data like integration. Based on my initial review of the data the series appeared to be non-stationary. I used the augmented Dickey-Fuller (ADF) test for the order of integration, I(0) or I(1). I am checking for deterministic or stochastic trends. The second step includes estimation of a Vector Autoregressive (VAR) model and the application of Johansen's maximum likelihood test to identify long-run relations among electricity consumption and possible determinants discussed before. The third step is to reduce the model through testing parameter restrictions and find a more parsimonious model. In addition, I conducted tests for various forms of exogeneity involving inference, forecasting, and possibly policy analysis.

A. Integration

The ADF unit root test is applied to each variable for the 1974-2006 period. The ADF test results are presented in Table III.2. The null hypothesis on the left half are whether the series have a unit root; on the right hand side are tests of the null hypothesis that the series are I(2). These are written level and first difference form. The ADF t-tests reveal that a unit root in the level cannot be rejected and that a unit root in the differences is rejected. Thus I conclude that the logarithm of variables for residential electricity consumption per person (*eh*), industrial electricity consumption per worker (*ep*), the residential electricity price (*ph*), the industrial electricity price *(pp)*, real disposable income per capita(*i*), real output

(*gdp*), and stock of household furnishings (*a*) are I(1) series. However based on the graphical analysis, it appears as if the *(pp)* is influenced by a one time shock. Thus I am uncertain if it should be treated as I(1) or I(0).

B. VAR lag selection

The following general VAR model is estimated for capturing the dynamics, testing for stability, and selecting the lag order:

$$y_t = c + A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + \varepsilon_t$$

where $y_t = (eh, i, ph, a)'$ or $y_t = (eh, i, a)'$ for residential electricity demand; and

$= (ep, gdp, pp)'$ or $y_t = (ep, gdp)'$ for industrial electricity demand;

A_i = the matrix of coefficients;

ε_t = the vector of error terms; and

c = constant terms.

Note that two specifications are examined for the residential sector and the industrial sector. The price of electricity has been removed in the second specification, because the results suggested that prices may not be part of the long-run determinants and would affect electricity demand only through short-run channels. This may have been part of economic and electricity policy. Prices were administered in a way so as not to deter economic development.

Table III.3 summarizes VAR lag selection results of the Akaike Information Criterion (AIC), and Schwarz Bayesian Criterion (SC). For two of the residential electricity demand models, AIC suggests lag three lags, while SC suggests only one lag. Thus, the results are inconclusive. I decided to keep two specifications for the successive analysis. For industrial electricity demand model, both AIC and SC suggest lag 1. However, this implies the zero lag or no short-run dynamics, when VAR cointegrating models are transformed into a vector error correction model (VECM), and therefore, is not of much interest. Accordingly, to keep at least one lag in VECM, we assumed two lags for the VAR model for the industrial sector. I use the two lag specification, but also retain the three lag version for residential electricity consumption.

A necessary condition for a credible Johansen cointegration test includes well-behaved error terms in VARs: the vector error terms ε_t should follow the white noise process. Diagnostic statistics of the estimated VAR residuals like autocorrelation, normality and heteroskedasticity are presented in Table III.4. All the models, except Model II, the residential electricity system with three lags appear to pass the tests at the 5% level. The null hypothesis of no-autocorrelation and normality can be rejected for Model II at the five percent significance level, suggesting that the estimated residuals may not be normally distributed and have a residual serial correlation problem.

C. Johansen cointegration test

With this preliminary data analysis in hand, I moved on to unrestricted Johansen cointegration tests (Table III.5). In testing for cointegrating relations or vectors I included deterministic variables like the *debt crisis dummy* (1984=1985=1) and a time trend to see if they provided significant explanatory power. None were deemed necessary in the residential models, but I did retain both variables in the industrial demand model. The trend term was only included in the cointegrating vector. The sample period for estimation remains the same as 1974-2006.

- **Residential electricity demand model:** The trace tests indicate weak evidence of the existence of a theory-consistent cointegrating equation, if four variables (*eh*, *i*, *ph*, *a*) are put into the cointegration space altogether. With two lags (Model I), the null hypothesis for no cointegrating is rejected at the 6 percent significance level (with the corresponding trace statistics, 47.371). However, the β coefficient for the price is -0.292 which implies a positive sign which does not make economic sense. With three lags (Model II), the trace test indicates the non-existence of a cointegrating relationship with the probability value of 0.137. Accordingly, at this stage, we drop these two models (Model I and II) from the subsequent analysis. If we drop the price from the cointegrating space, the trace tests indicate relatively stronger evidence of the existence of cointegrating equation (Model III and IV), with the null hypothesis of *r=0* rejected at less than one percent significance levels. All β coefficients show the expected signs and are significant. The α coefficients in the *d(eh)* equations for both models are negative, while those in *d(i)* equations are

positive, supporting the expected error correction mechanism. However, the α coefficients in *d(a)* equations in both models are negative. The sign here appears to make economic sense. If electricity consumption is above the long-run equilibrium and expected to slow down, this would suggest that stock of electric appliances may slow as well.

- **Industrial electricity demand model:** The industrial electricity price did not enter the cointegrating relation in the industrial electricity model. When it is removed the trace statistic is highly significant (2%) with the null hypothesis of no cointegration, and all β and α coefficients present the expected signs (Model V). For the industrial sector, a time trend also enters into the cointegrating vector.

Next, I examined several parameter restriction tests on the cointegrating coefficients β *and* α . These are presented in Table III.6.

- First, the results from both cointegrating systems suggested that electricity prices did not enter the equilibrium relations. The discussion below focuses on models III and IV for residential electricity and model V for industrial electricity.

- Second, as discussed before, the negative α coefficients in the *d(a)* equations imply that a positive disequilibrium in the cointegration relationship would suggest that growth in the stock of appliances *a* would slow down. I was skeptical of this result

empirically. Since the speed of adjustment coefficients are numerically small, I test the null hypothesis of *α coefficient to d(a)* =0. For Model III (two lags in the original VAR model), the null hypothesis can be rejected at a relatively high 4.3 percent significance level, suggesting that *d(a)* =0 may not hold. While for Model IV (three lags in original VAR), it cannot be rejected; the p-value is 33.7 percent implying the validity of the hypothesis. I will continue with both specifications for now.

- Third, one interesting hypothesis is income elasticity of residential electricity is unity or β coefficient on income is equal to one. The Lagrangean Likelihood test results show that the joint null hypothesis of *α coefficient to d(a)* =0 and *β coefficient to i* =1 cannot be rejected at a five percent significance level for both Model III and IV. This has a rather strong implication: a long-run elasticity of residential electricity demand with respect to income is equal to one.

- Finally, I test for the weak exogeneity of income, which relates to two long-lasting question posed by many researchers in the previous studies: whether there is causality between consumption and income and if so which does the causality run, from electricity consumption to income, vice versa, or is there feedback? The Lagrangean Likelihood test results show that the joint null hypothesis of *α coefficient to d(a)* =0, *β coefficient to i* =1, and *α coefficient to d(i)* =1 can be rejected at a less than one percent significance level for both Model III and IV

respectively. This suggest that income is not weakly exogenous, and there is feedback in Granger causality between electricity consumption and income.

Until now, the statistical tests do not show clearly which VAR model, with two or three lags, best represents the data generating process, because both models broadly passed the diagnostic, cointegrating, and parameter restriction tests. Note that parameter estimates for both equations are quite similar. I proceed with Model IV (three lags in the original VAR model) as it nests Model III and impose the parameter restrictions on the *α coefficient to d(a)* =0 and *β coefficient to i* =1, as the long-run cointegrating relation for residential electricity demand. The Lagrangean Likelihood test results support these parameter restrictions more strongly than Model III. In sum, the final long-run residential electricity consumption model can be summarized as:

$$eh = i + \underset{[7.023]^{**}}{0.257} \times a \qquad \text{(3 or Model IV)}$$

$$\text{with speed of adjustment coefficients } \alpha: \begin{cases} d(eh)equation & -0.320[-2.775]^{**} \\ d(i)\,equation & 0.218\,[2.940]^{**} \end{cases}$$

where [ ] indicates t statistics and $d(x)$ indicates the first difference of the variable x.

For the industrial sector model there does not appear to be a unit production or gdp per worker elasticity; the coefficient is well below unity and the standard error is relatively small. I consider a single parameter restriction: testing for weak erogeneity of *gdp*. The Lagrangean Likelihood tests suggest that the joint null hypothesis of *α coefficient to d(gdp)* =0 can be rejected at the one percent significance level (Model V in Table III.6). Thus, I

find results similar to the previous residential electricity demand model. There is feedback between industrial electricity demand and *gdp;* they appear to Granger-cause each other. Because of rejection of this parameter restriction test, we keep the earlier unrestricted model (Model V in Table III.5) as our final long run model for production sector electricity demand model. Note there is a small 0.4% per year increase in electricity consumption per worker.

$$ep = \underset{[2.743]^{**}}{0.299} \times gdp - \underset{[-3.408]^{**}}{0.004} \times t \qquad \text{(4 or Model V)}$$

$$\text{with speed of adjustment coefficients } \alpha: \begin{cases} d(ep)equation & -0.437[-2.517]^{**} \\ d(gdp)equation & 0.214\,[2.623]^{**} \end{cases}$$

where [ ] indicates t statistics and $d(x)$ indicates the first difference of the variable x.

The Philippines is a developing country and emerging market country. Economic activity has been subject to several internal shocks, external shocks, and structural changes. A natural test would be whether the models are stable and whether the estimated parameters appear constant. The models with no parameter restrictions (in Table III.5) can be re-estimated recursively, starting with the initial sample of up to 1990, and adding yearly observations through 2006. Figure III.5 presents *β coefficients to i* and *a* for residential Model IV and *β coefficients to gdp* for industrial Model V. All coefficients appear reasonably stable. However, the band of two standard errors is relatively large, suggesting that the precision of those *β coefficients* could be in question. We will come back to this issue when we apply the model for forecasting of electricity demand in the Philippines in Section VII.

VI. ELECTRICITY DEMAND MODELS

This section presents the results from the conditional vector error correction models. For each of the residential sector and industrial sector electricity demand models I map the models in levels to first differences and include the lagged deviation in cointegrating relation, plus additional variables which may have short-run effects. The residential sector model is in terms of the change in per capita consumption and the change in income per capita; the system is conditioned on electricity prices and household appliances. The industrial sector model is in terms of the change electricity consumption per worker and the change in real gdp per worker; the system is conditioned on electricity prices. The vector error correction models (VECM) combine the above long-run relationships with short-run dynamics in fully-specified electricity demand models. I use the modeling strategy of a general-to-specific approach advocated by Hendry and others (for example, see Hendry, 2000).

I begin by estimating general unrestricted models in each sector. In the residential electricity demand model, the short-run variables include *deh(lag)*, *di(lag)*, *da(lag)*, *dph(lag),* where lag takes 1 to 2, while for the industrial sector electricity demand model, the variables include *dep(lag)*, *dgdp(lag)*, *ppgap(lag)*, where lag takes 1. The selection of the initial lag order reflects the results of VAR analysis in the previous section. I include a term *ppgap* for the deviation of industrial prices around their mean in the industrial sector electricity demand model, instead of the first difference as noted in the ADF test results. I was uncertain how to include this variable because industrial sector electricity prices could

be *I(0)* and not *I(1)*. However, to make the interpretation of VECM representation easier, a difference of the level of pp and its mean (1974-2006) average, *ppgap*, is used. This implies that expectation of *ppgap* is zero, and thus, at the static equilibrium (none of the variables changes), *ppgap* is zero. Furthermore, a debt crisis dummy (1984=1985=1), temperature (*temp*), and constant are included as additional explanatory variables.

A. Residential sector

The unrestricted bivariate model for the residential sector is presented below and found in Table III.7 – Model R1.

$$deh = \underset{[-2.44]^{**}}{-0.29} \times LRR(-1) - \underset{[-0.28]}{0.05} \times deh(-1) + \underset{[0.99]}{0.19} \times deh(-2) - \underset{[-0.08]}{0.02} \times di(-1) - \underset{[-1.08]}{0.21} \times di(-2)$$
$$- \underset{[-0.43]}{0.32} \times da(-1) - \underset{[-0.68]}{0.63} \times da(-2) - \underset{[-2.08]^{*}}{0.31} \times dph(-1) - \underset{[-0.26]}{0.05} \times dph(-2) + \underset{[0.03]}{0.19} + \underset{[0.56]}{0.04} \times debtdummy$$
$$- \underset{[-0.14]}{0.01} \times temp$$

$$di = \underset{[3.12]^{**}}{0.24} \times LRR(-1) - \underset{[-1.44]}{0.18} \times deh(-1) + \underset{[0.15]}{0.02} \times deh(-2) - \underset{[-2.67]^{**}}{0.46} \times di(-1) - \underset{[-3.41]^{***}}{0.43} \times di(-2)$$
$$+ \underset{[1.50]}{0.71} \times da(-1) - \underset{[-1.41]}{0.85} \times da(-2) - \underset{[-1.52]}{0.15} \times dph(-1) + \underset{[0.68]}{0.08} \times dph(-2) + \underset{[0.56]}{0.23} - \underset{[-6.44]^{***}}{0.28} \times debtdummy$$
$$- \underset{[-0.46]}{0.03} \times temp$$

where *LRR* is the long-run cointegration relationship (the error correction term) lagged one period in the residential sector (equation 3), and [ ] indicates t statistics.

The coefficient to the long-run relationship, *LRR*, is negative and positive in the residential and income equations respectively. Only one other variable is significant in the *deh* equation out of ten. Four variables are significant in the *di* out of eleven. It appears as though the model is over-parameterized. I conduct a process of dropping insignificant variables to reduce the model. I place zero restrictions on those pairs of variables with the

least significant level one by one from the system altogether and re-estimate the model by OLS.

This produces seven models; see Table III.7, Model R2-R7. These models are nested in the general unrestricted model. I test for loss of explanatory power using F test statistics. The model reduction test results are summarized in Table III.8. F tests cannot reject the null hypothesis against the alternatives for sequential model reductions from Model R2 to Model R6. The probability value of rejection from Model R6 to Mode R7 is five percent, much smaller than the previous stage. However, I selected Model R7 as the most simplified and parsimonious model at this stage. In the process, the lagged change in household appliances from two years ago is dropped. The sign was significant but negative in the *deh* and insignificant in the *di* equation. There was little change in the other estimated parameters.

$$deh = \underset{[-2.22]^{**}}{-0.22} \times LRR(-1) - \underset{[0.01]}{0.00} \times di(-1) - \underset{[-0.69]}{0.12} \times di(-2) - \underset{[-2.45]^{**}}{0.33} \times dph(-1) - \underset{[-2.14]^{**}}{0.92} + \underset{[0.84]}{0.06} \times debtdummy$$

$$di = \underset{[3.60]^{***}}{0.23} \times LRR(-1) - \underset{[-4.81]^{***}}{0.66} \times di(-1) - \underset{[-3.93]^{***}}{0.44} \times di(-2) - \underset{[-1.62]}{0.14} \times dph(-1) + \underset{[3.75]^{***}}{1.01} - \underset{[-6.56^{***}}{0.28} \times debtdummy$$

where [ ] indicates t statistics.

Sample period=1974-2006; R^2=0.74; Vector AR 1-2 test: F(8,44)=1.3740 [0.2344]; Vector Normality test: Chi^2(4) =4.7442 [0.3146]; Vector hetero test: F(18,51)=0.56981 [0.9048]; Vector hetero-X test: F(27,44)=0.41130 [0.9919].

where [ ] indicates probability value.

Coefficients to the long-run relations have changed little, and short-run variables are significant in the simplified model compared to the original model (Model R1). However, there remain insignificant short-run and dummy variables in both equations. Accordingly, we step further and drop insignificant variables in each equation, and re-estimate the system. I need to pay attention that the right-hand-side of the system is no longer identical, and that OLS is inconsistent in a simultaneous system. Hence, I use the Full Information Maximum Likelihood (FIML) method. The final estimated VECM for residential electricity is:

$$deh = \underset{[-1.89]*}{-0.17} \times LRR(-1) - \underset{[-2.36]**}{0.27} \times dph(-1) - \underset{[-1.81]*}{0.69}$$

$$di = \underset{[3.82]***}{0.24} \times LRR(-1) - \underset{[-5.75]***}{0.72} \times di(-1) - \underset{[-3.56]***}{0.37} \times di(-2) + \underset{[3.97]***}{1.08} - \underset{[-7.20]***}{0.29} \times debtdummy \quad \text{(5 or Model R8)}$$

where [ ] indicates t statistics.

Sample period=1974-2006; Vector AR 1-2 test: F(8,48)=1.5235 [0.1741]; Vector Normality test: Chi^2(4) =8.2299 [0.0835]; Vector hetero test: F(27,50)=0.5123 [0.9682]; Vector hetero-X test: F(48,30)=0.7063 [0.8614].

where [ ] indicates probability value.

The estimated bi-variable model (5 or Model R8 in Table III.7) is highly parsimonious. The coefficient to the long-run relation has fallen numerically (-0.29 to -0.17) but not statistically for the *deh* equation from the original model (Model R1), while has changed little for *di* equation. All short-run variables are highly significant and drastically simplified. All diagnostic test results are satisfactory, and Chi-square statistics support a reduction of the model from R7 to R8. Parameter stability test suggest that one step

residuals broadly stay within two standard deviation bands (95 percent confidence), while break-point F test statistics (Chow test) remain below the one percent line. It all indicates non-rejection of constancy as shown in Figure III.6.

B. Industrial sector

The unrestricted bivariate model for the industrial sector is presented below and found in Table III.9, Model P1.

$$dep = \underset{[-2.37]**}{-0.43} \times LRP(-1) - \underset{[-0.54]}{0.09} \times dep(-1) + \underset{[0.99]}{0.30} \times dgdp(-1) - \underset{[-1.21]}{0.11} \times ppgap - \underset{[-0.47]}{0.03} \times pp(-1)$$
$$+ \underset{[2.37]}{1.68} + \underset{[0.17]}{0.01} \times debtdummy + \underset{[0.44]}{0.02} \times temp$$

$$dgdp = \underset{[2.17]**}{0.19} \times LRP(-1) - \underset{[-0.57]}{0.05} \times dep(-1) - \underset{[-0.91]}{0.13} \times dgdp(-1) - \underset{[-1.14]}{0.05} \times ppgap - \underset{[-0.26]}{0.09} \times pp(-1)$$
$$- \underset{[-2.15]}{0.73} - \underset{[-5.22]***}{0.13} \times debtdummy - \underset{[-0.52]}{0.01} \times temp$$

where *LRP* is the long-run cointegration relationship (the error correction term) in the residential sector (equation 4), and [ ] indicates t statistics.

The coefficient to the long-run relationship, *LRP*, is significant and of the expected sign. Once again the model appears to be over-parameterized. I follow the same procedure used in the residential reduction tests. I re-estimate the model by OLS, and obtain four additional models (Table III.9, Model P2-P5). The F tests for sequential model reduction are summarized in Table III.10, which shows no rejection of the null hypothesis in the sequence from Model P1 to Model P4. The null hypothesis of the validity of model reductions from Model P4 to Model P5 cannot also be rejected, but is marginal at six percent. The one year lag change in *gdp* is removed between Model 4 and Model P5. The

two coefficients were not significant in the former and it was negative in the *dgdp* equation. Model 5 is presented below:

$$dep = \underset{[-2.79]^{***}}{-0.38} \times LRP(-1) - \underset{[-2.39]^{***}}{0.15} \times ppgap + \underset{[2.79]^{**}}{1.46} + \underset{[0.93]}{0.004} \times debtdummy$$

$$dgdp = \underset{[2.48]^{**}}{0.16} \times LRP(-1) - \underset{[-1.28]}{0.04} \times ppgap - \underset{[-2.44]}{0.61} - \underset{[-5.51]^{***}}{0.12} \times debtdummy$$

where [ ] indicates t statistics.

Sample period=1974-2006; R^2=0.84; Vector AR 1-2 test: F(8, 48)=1.0780 [0.3943]; Vector Normality test: Chi^2(4) =5.9026 [0.2065]; Vector hetero test: F(15, 58)=0.92149 [0.5460]; Vector hetero-X test: F(21, 55)=0.79580 [0.7122].

where [ ] indicates probability value.

Most coefficients are significant, but not all. Thus, I employ FIML in the final step to further simplify the system. The result is:

$$dep = \underset{[-2.70]^{**}}{-0.36} \times LRR(-1) - \underset{[-2.71]^{**}}{0.12} \times ppgap + \underset{[2.71]^{**}}{1.38}$$

$$dgdp = \underset{[2.99]^{**}}{0.19} \times LRR(-1) - \underset{[-2.94]^{***}}{0.71} - \underset{[-8.37]^{***}}{0.13} \times debtdummy \qquad \text{(6 or Model P6)}$$

where [ ] indicates t statistics.

Sample period=1974-2006; Vector AR 1-2 test: F(8, 50)=1.7419 [0.1117]; Vector Normality test: Chi^2(4) =8.6153 [0.0715]; Vector hetero test: F(15, 61)=0.9150 [0.5524]; Vector hetero-X test: F(21, 57)=0.8606 [0.6377].

where [ ] indicates probability value.

Coefficients to the long-run relations have changed little, and all short-run variables are significant in both *deh* and *dgdp* equations. This final model (6 or Model P6 in Table III.9) satisfies all diagnostic tests. Various parameter constancy tests broadly confirm stability of the model, except year 2001, as evidenced by the fact that one-step residuals lie within 95 percent confidence interval (±2 standard deviations), while break-point Chow test statistics are below one percent significance level (Figure III.7). In year 2001, the *dgdp* equation suggests some evidence of parameter instability. However, this could be largely attributed to shocks caused by political crises, following the outbreak of the President Estrada corruption scandal, and resulting temporary economic slump.

VII. Further Interpretation and Discussion of the Results

Below I summarize the main findings on the residential sector and industrial sector electricity consumption in the Philippines.

- **Integration and Cointegration:** There is evidence in favor of all variables following an I(1) process except one. Residential electricity consumption, income, and stock of appliances have a long-run cointegration relationship. However, such a relationship cannot be found if the residential electricity price is included. In the industrial sector, industrial electricity consumption and GDP are cointegrated. The industrial electricity price did not appear the cointegration space. However, it did enter the VECM.

- **Long-run income elasticities:** The long run income elasticity of residential electricity demand is one, while that of the stock of electric appliances is 0.257. The estimated income elasticity is similar to the ones found (see Table III.1) for Cyprus (Zachariadis and Pashourtidou, 2007) and Taiwan (Holtedahl and Joutz, 2004). For industrial electricity demand, the income elasticity is 0.299, much lower than that for residential electricity demand.

- **Short-run income elasticities:** For both residential and industrial electricity demand models, the short-run income variables, *di(lag)* or *dgdp(lag)*, are found insignificant, implying zero short-run income elasticity. This is consistent with results of other studies that the income elasticity is generally smaller in the short-run than in the long-run (see for example, Silk and Joutz, 1997).

- **Electricity price:** Electricity prices are important determinant of short-run electricity demand for both residential and industrial sectors. In the residential electricity demand model, a change in the residential electricity price affects a change in electricity demand, with the short-run price elasticity estimated as -0.27. This is similar to Australia and U.S. cases (Narayan et. al., 2005, and Silk and Joutz, 1997), but higher than other countries, where data are available, such as Denmark, New Zealand, Taiwan, and Turkey. In the industrial electricity demand model, interestingly, it is the level of the electricity price, rather than the change, that affects demand as a short-run factor.

- **Granger causality:** For both residential and industrial electricity demand functions, there is no evidence that income variables are weakly exogenous. This implies that electricity demand and income (or GDP) cause each other in the Granger sense.

- **Temperature:** In other countries, such as Denmark (Bentzen and Engsted, 1992), U.S. (Silk and Joutz, 1997), Greece (Hondroyiannis, 2007), and Namibia (De Vita, et. al., 2005), temperature is an important determinant of electricity demand. However, in the case of the Philippines, this variable is insignificant, and further research would be needed.

Policy implications

My empirical models provide a tool for policy makers. I compare projections for future electricity demand based on my estimated models and the government's official one. The latest official forecasts of power demand are Power Development Plan (PDP) 2005-2014 by the Philippines' Department of Energy. With the assumption of an average GDP growth of 6.7 percent for 2005-2014, PDP envisages total electricity demand to grow 7.6 percent on average during these periods. This implies that the elasticity of electricity demand respect to GDP is 1.134, which is much greater than our estimated results based on the long-run cointegration model (for the residential sector, the elasticity is 1.0, and for the industrial sector, 0.299).

As discussed before, estimated long-run β coefficients are reasonably stable (Figure III.5), but the two-standard error bands are relatively large. Accordingly, forecasting can take account of the degree of imprecision of the estimated parameters. We forecast total electricity demand assuming that (i) income and GDP growth will average 6.7 percent a year (consistent with the government's plan); (ii) the stock of electronic appliances will grow at an annual rate of 6.7 percent (for simplicity, the same growth rate as income and GDP); and (iii) population and labor will grow 2 percent a year (the government's plan). Figure III.8 plots four forecasting scenarios for total electricity demand, and these are based on: the government's PDP (Scenario I); unrestricted long-run models for residential and industrial production (Model IV and V in Table III.5: Scenario II); the same model used in (II) but with β coefficients to income and GDP $+2\sigma$ (Scenario III); and the same model used in (III) but with β coefficients to income and GDP -2σ (Scenario IV) With this set up, the area between Scenario III and IV can be interpreted as 95 percent confidence region.

The Figure III.8 clearly shows that the government's PDP forecast lies on the upper border of the 95 percent confidence region, suggesting that according to our estimated model, the government's forecast may have over-predicted future electricity demand. This is consistent with observations of Figure III.8. The government's electricity consumption forecast appears to have an accelerated trend compared with the historical trend (the average annual electricity consumption growth rate was 6 percent for 1993-2006, below the government's forecasted growth of 7.6 percent annually). Furthermore, such a concern about over-prediction would be warranted if income and GDP grow lower than their

predictions assumed for forecasting electricity consumption. In addition, the excess capacity would put a financial burden on the government. The burden is not solely for the financing the electric generating capacity. These funds may have been used for alternative projects with relatively high returns.

Another important implication that can be drawn from my empirical models regards electricity pricing policy. This is particularly so, because in the Philippines, the setting of electricity tariffs is a contentious and sensitive social issue. It has often caused public strife. Moreover, concerns about adverse impact on industry competitiveness arising from higher tariffs have always been high, because the electricity tariffs for industrial users in the Philippines were among the highest in the ASEAN countries. With increased competition from China and other neighboring countries, higher electricity prices would undermine the Philippines' attractiveness to foreign investors. Our models suggest that consistent with those concerns, a higher level of the electricity price in the industrial sector would reduce both industrial electricity demand and GDP, although the latter is not statistically highly significant (Model P5 in Table III.9).

Note that the impact of higher prices on electricity demand is potentially very large, the extent of which depends on the level of price gap, which is a degree of divergence of the electricity price level from its mean. To verify this point, assuming that the economy is in equilibrium and thus error correction terms are zero in *dep* equation in (6) or Model P6,

$dep = -0.12 \times ppgap + 1.38$.

In 2006, an industrial electricity price was 160 centavos (in the 1985 price), and the average electricity price for the 1974-2006 periods was 141 centavos. Thus, the 2006 price was higher than the 1974-2006 average by 13 percent (in logarithm, approximately *pp*=0.13). This would lead to a 1.6 percent reduction (*dep*=-0.12×0.13 =-0.016) in industrial electricity demand in the short-run. This is inelastic and commonly found in the literature. See the references in Table III.1. However, the larger the price gap, *ppgap*, the more elastic or greater a drop in electricity consumption demanded would be. For example, for the electricity price 20 percent above the average, electricity consumption demanded would fall 2.5 percent, for 50 percent above, a 5 percent decline, and for twice above, a 8 percent decline (Figure III.9).

The bottom line indicates that a sharp change in the electricity price would potentially cause a significant impact on electricity demand in the industrial sector. GDP would also be affected through the error correction mechanism. If industrial electricity demand declines 8 percent (corresponding to the price twice above the average), GDP would fall in the subsequent period 1.6 percent (= –0.08×0.2).

VIII. CONCLUSIONS

My research investigates annual electricity consumption for the Philippines. For the residential sector, a long-run cointegrating relationship is found among residential electricity consumption, income, and stock of electric appliances, while for the industrial

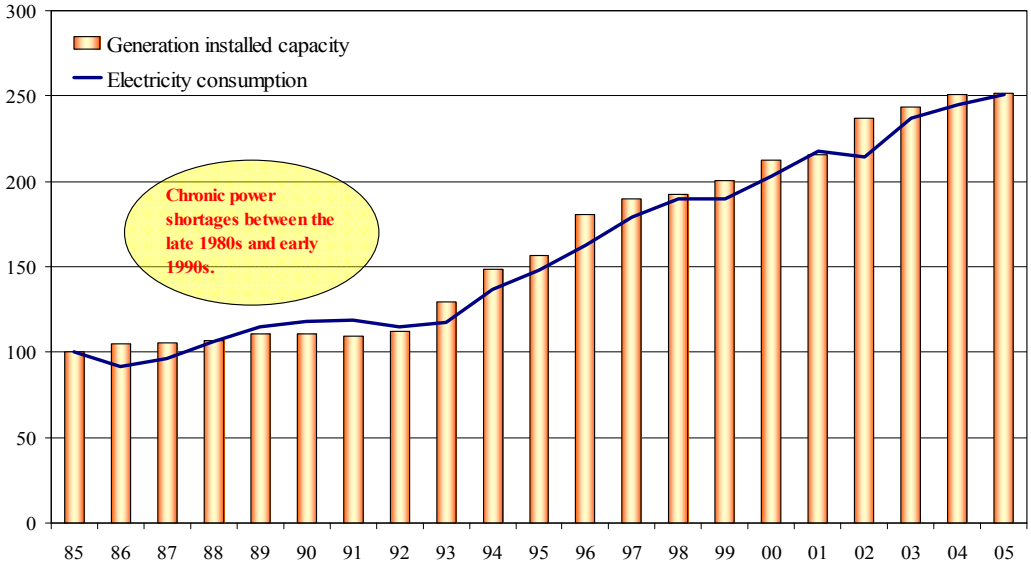
sector, a long-run relationship holds for industrial electricity consumption and GDP. A short-run dynamic is relatively simple for both sectors, and we found that consistent with the other studies in the literature, the long-run elasticities with respect to income or GDP are larger than the short-run elasticities. Also, I found that the electricity price was an important determinant of electricity demand in the short-run, but not in the long-run. Granger causality tests also verified that in the Philippines, the direction of causality in the residential and industrial electricity demand models runs both ways: electricity and income (or GDP) granger-cause mutually.

Despite the small sample size, estimated models have broadly passed various specification and stability tests. However, a band that estimated parameters, such as β coefficients to income and GDP in the long-run models, can take is relatively wide, which should be appropriately taken account when we use the models for the forecasting purpose. The prediction of the level of total electricity consumption based in the estimated model is lower than the government's forecast in the Philippine PDP. In fact, the government's forecast lies in the upper bound of the forecast range suggested by our model. Finally, a sharp upward deviation of industrial electricity price from its mean would cause a large decline in electricity consumption demanded in the short-run. Such decline in the electricity consumption could also affect GDP to the extent that depends on the size of deviation through the error correction mechanism.

Further improvement in data quality would help strengthen statistical analysis. The availability of higher frequency data could broaden the scope for the time series analysis.

Data for the stock of appliances could be officially compiled, because the appliances consist of one of the important determinants of residential electricity demand, as shown in my research.

Figure III.1. Generation Installed Capacity and Electricity Consumption (1985=100) 1/



1/ Underlying data: Megawatt for generation installed capacity and kWh for electricity consumption.

Figure III.2. Cooling Degree Days

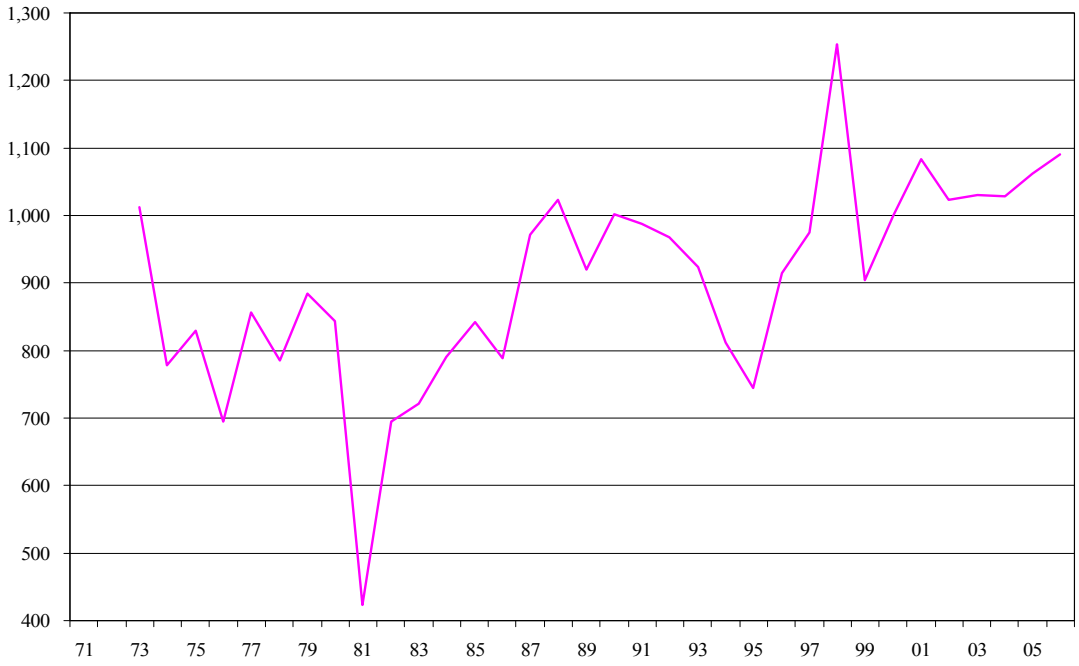


Figure III.3. Residential Electricity Consumption Per Capita
(In logarithm)

Figure III.3a. With Income Per Capita (i)

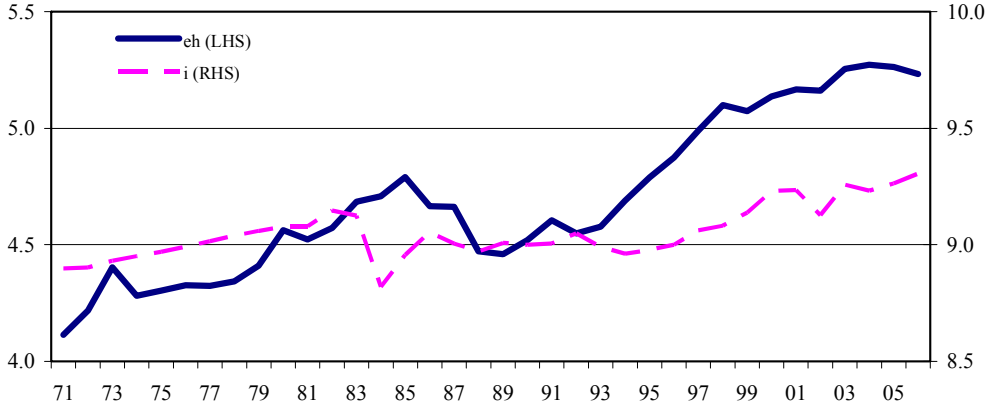


Figure III.3b. With Residential Prices (ph)

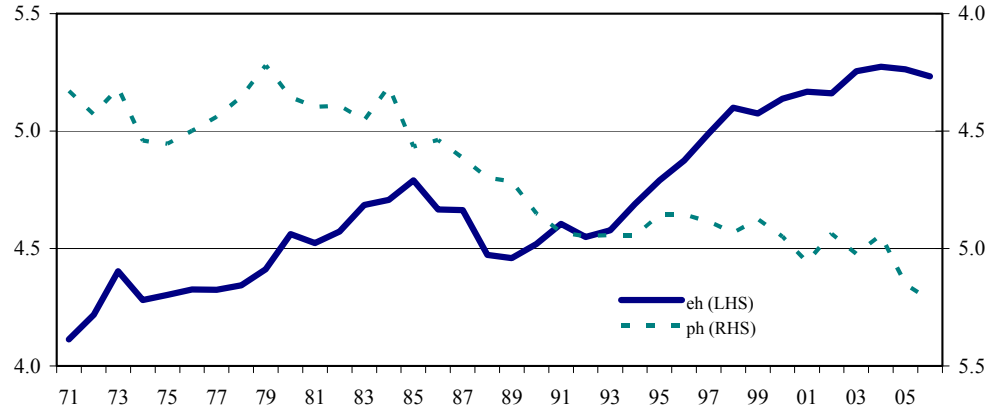


Figure III.3c. With Stock of Electronic Appliances Per Capita (a)

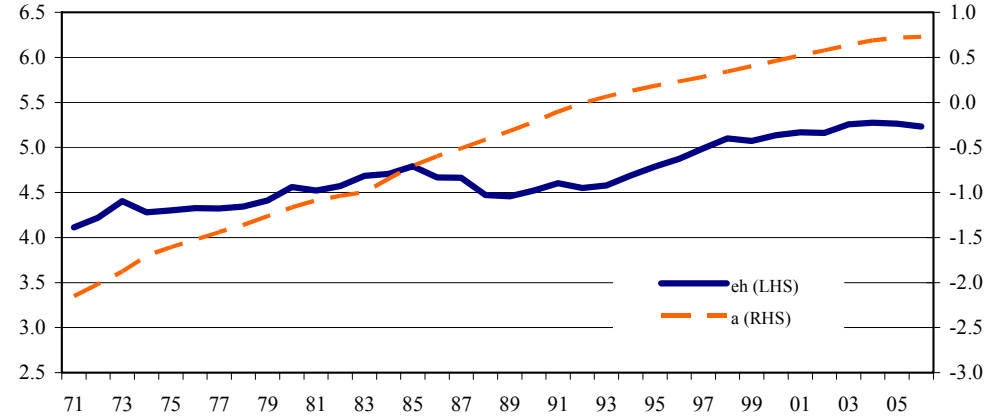


Figure III.4. Industrial Electricity Consumption Per Labor
(In logarithm)

Figure III.4a. With GDP Per Labor (gdp)

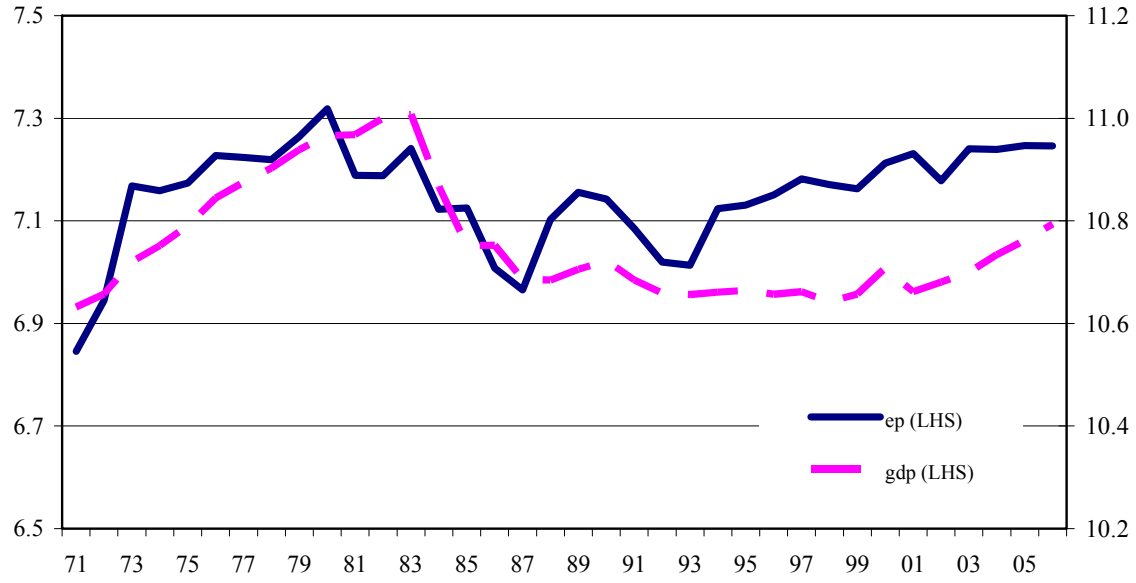


Figure III.4b. With Industrial Prices (pp)

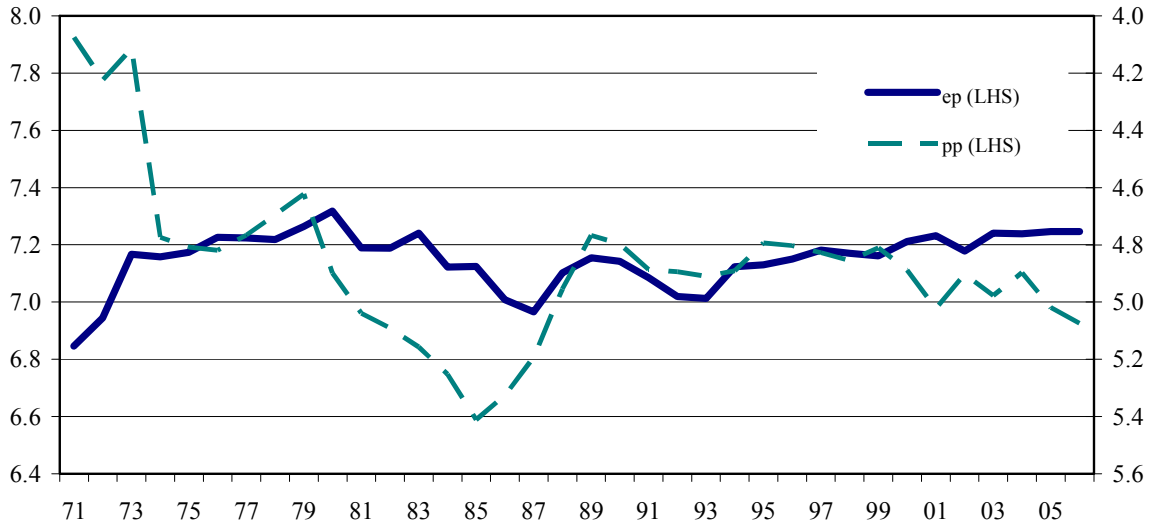
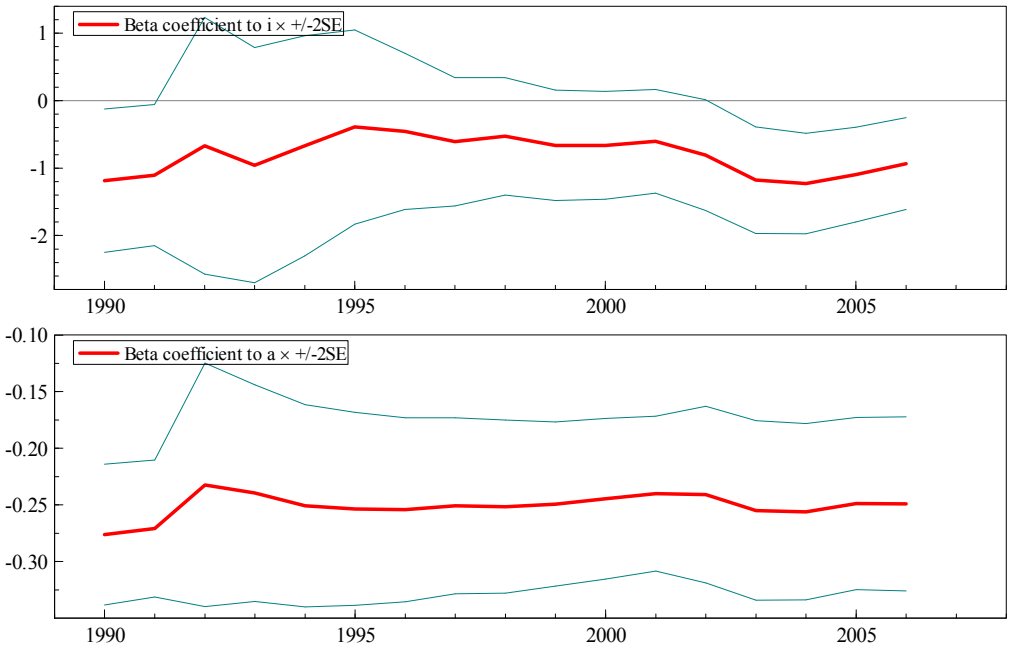


Figure 6. Production Sector Electricity Consumption Per Capita
(In the first difference of logarithm)

Figure III.5. Parameter Constancy by Recursive Regression

Residential electricity demand model (Model IV in Table 5.)



Industrial electricity demand model (Model V in Table 5.)

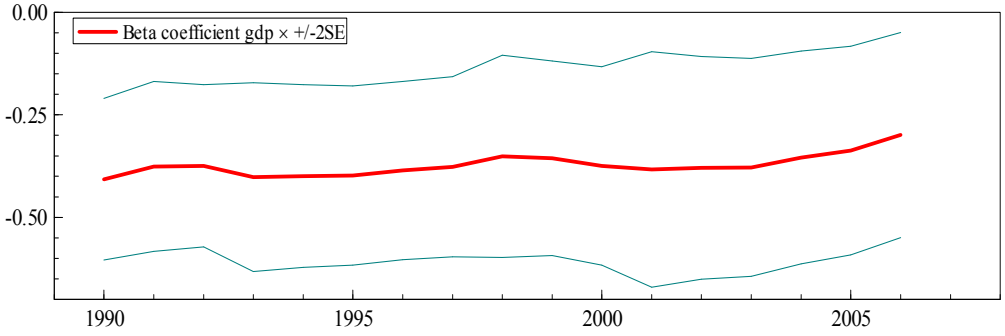


Figure III.6. Residential electricity demand model
Recursive Constancy Statistics: 1-step Residuals (r), One Step Chow (1 up), and Break Point Chow (Ndn) Tests

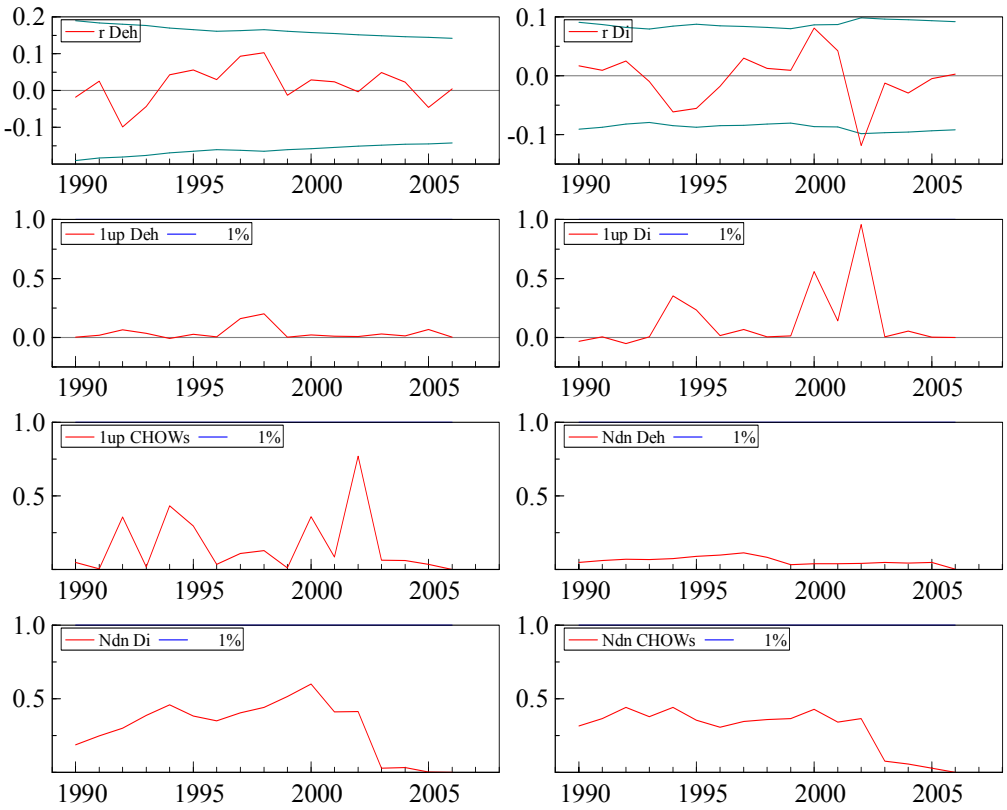


Figure III.7. Industrial Electricity Demand Model
Recursive Constancy Statistics: 1-step Residuals (r), One Step Chow (1 up), and Break Point Chow (Ndn) Tests

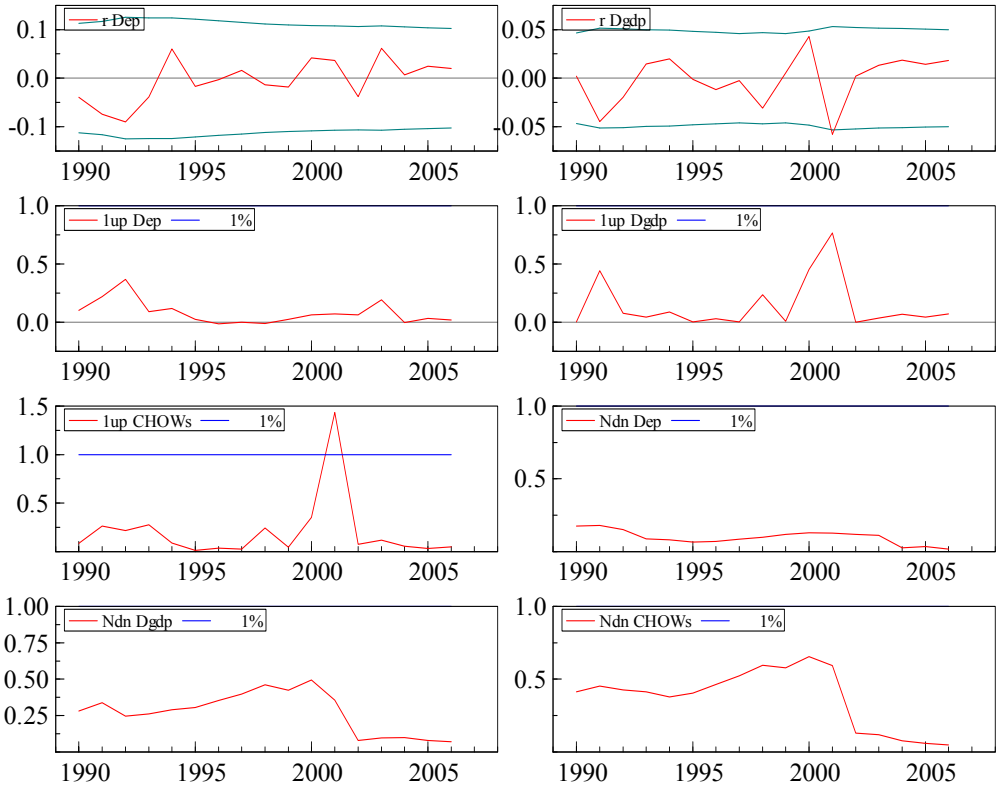


Figure III.8. Alternative Scenarios
(Total electricity demand; in gWh)

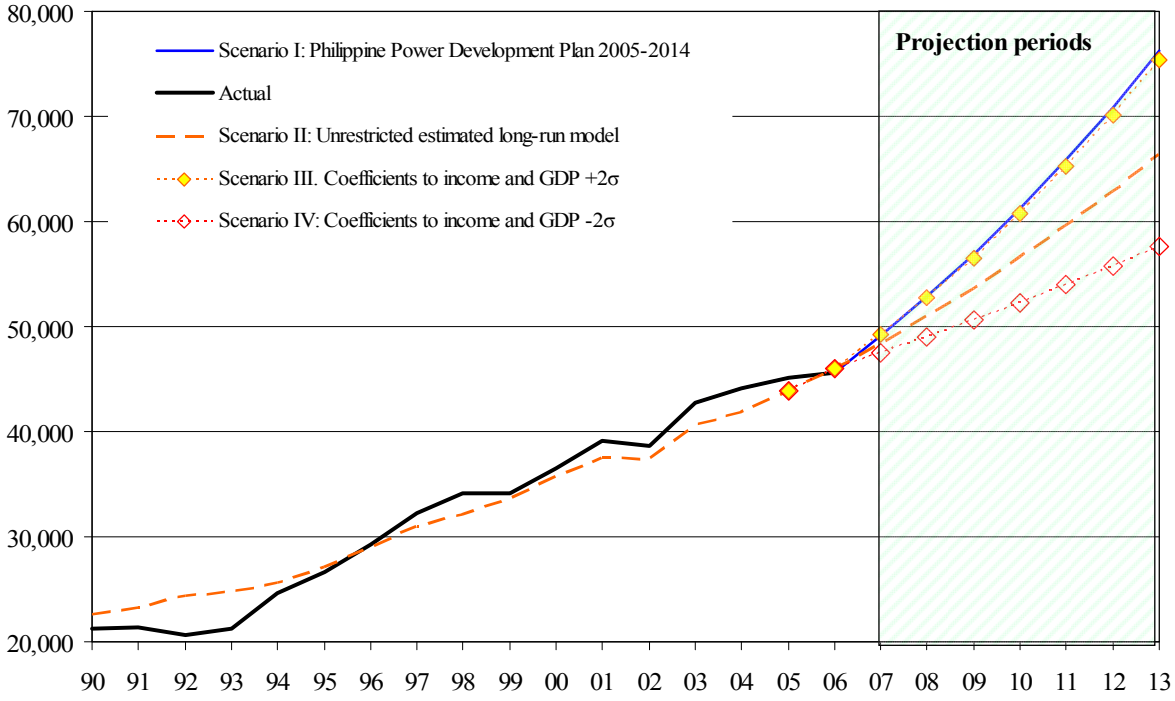


Figure III.9. Industrial Electricity Demand and Price

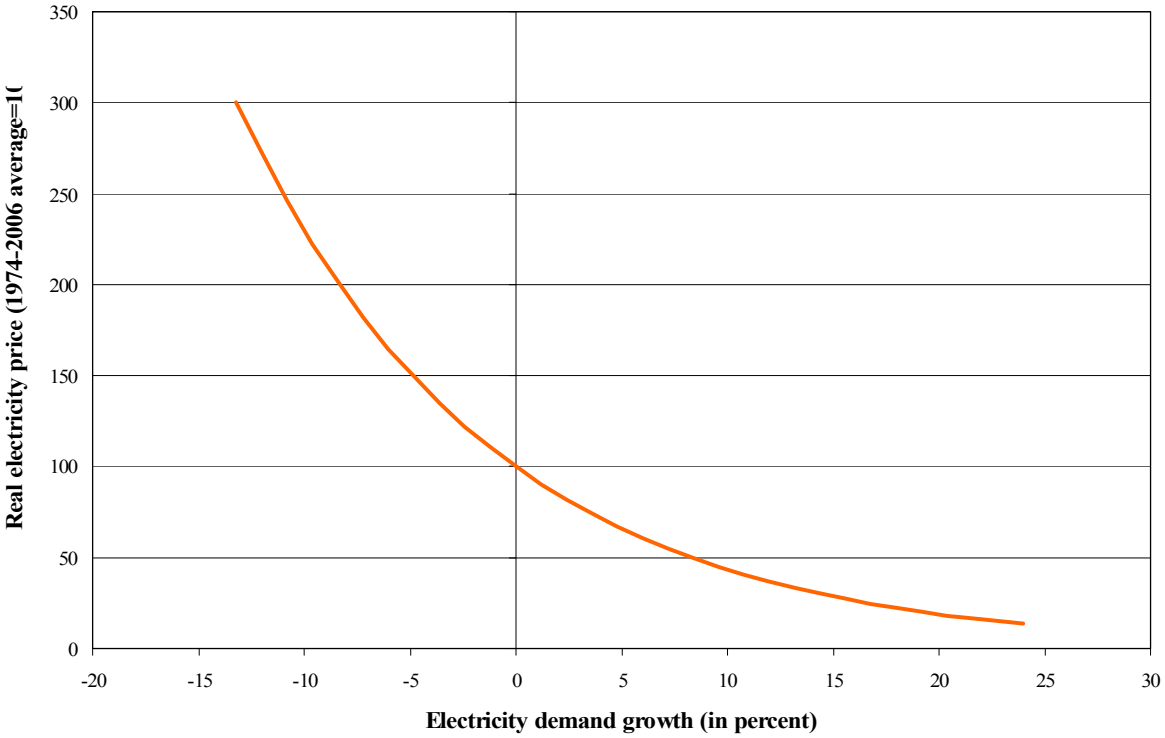


Table III.1. Summary of Selected Empirical Studies 1/

Country	Period	Dependent variable	Long-run model				Short run model					Source
			Income	Price	Others		ΔElectricity	ΔIncome	ΔPrice	Others		
Advanced countries												
Australia	1969-2000	Residential electricity	0.32	-0.54	1.69 0.16	Temperature Gas price	---	n.s.	-0.26	n.s. n.s.	ΔTemperature ΔGas price	Narayan et.a. (2005)
	1969q301998q2	Residential electricity	0.52	-0.95	0.21 0.38	Gas price Other fuel price	---	---	---	---		Akmal and Stern (2001)
Denmark	1948-1990	Total electricity	1.21	-0.47	---		n.s.	0.67	-0.14	0.41	ΔTemperature	Bentzen and Engsted (1992)
New Zealand	1960-1999	Total electricity	0.81	-0.59	0.42	Consumer price index	n.s.	0.24	-0.18	0.25	ΔConsumer price index	Fatai, et.al. (2003)
U.S.	1949-1993	Residential electricity	0.52	-0.48	0.26 0.16 0.04 0.06	Cooling degree days Heating degree days Interest rate Oil price	0.53	0.39	-0.28	0.16 -0.04 -0.01 0.11	ΔCooling degree days ΔHeating degree days ΔInterest rate ΔOil price	Silk and Joutz (1997)
Developing countries												
China	1952-2000	Total electricity	1.10	-1.32	---		---	---	---	---		Cheung et.al. (2003)
	1978-2001	Total electricity	0.78	-0.02	0.57 0.53 0.33	Population Heavy industry share in GDP Electricity consumption efficiency	0.09	0.76	0.02	-0.18 -0.48 -0.83	ΔPopulation ΔHeavy industry share in GDP ΔElectricity consumption efficiency	
Cyprus	1960-2004	Residential electricity	1.18	-0.43	---		n.s.	n.s.	n.s.	0.21	ΔTemperature	Zachariadis et.al. (2007)
		Commercial electricity	1.12	-0.30	---		n.s.	n.s.	n.s.	0.08	ΔTemperature	Zachariadis et.al. (2007)
Greece	1986m1-1999m12	Residential electricity	1.56	-0.41	-0.19	Temperature	0.24	0.20	---	---		Hondroyiannis (2004)
Israel	1973q1-1994q4	Residential electricity	---	-0.58	1.09	Private consumption	---	---	---	---		Beenstock et.al. (1997)
		Industrial electricity	0.99	-0.44	---		---	---	---	---		Beenstock et.al. (1997)
Namibia	1980q1-2002q4	Total electricity	0.59	-0.30	-0.36	Temperature	---	---	---	---		De Vita, et. Al (2005)
Taiwan	1957-1995	Residential electricity	1.04	-0.16	-0.16 3.91	Oil price Urbanization rate	---	0.23	-0.15	1.61 0.03	ΔUrbanization rate Temperature	Holtedahl and Joutz (2004)
Turkey	1968-2005	Residential electricity	0.70	-0.52	0.04	Urbanization rate	---	n.s.	-0.10	0.04	ΔUrbanization rate	Halicioglu (2006)
Philippines	1974-2006	Residential electricity	1.00	n.s.	0.26	Real stock of electronic appliances	n.s.	n.s.	-0.27	---		Ishi (2008)
	1074-2006	Industrial electricity	0.30	n.s.	---		n.s.	n.s.	---	-0.12	Real price gap from its historical average	Ishi (2008)

1/ "---" means that a variable is not included in the regression model, while "n.s." means that a variable is included in the regression model but is found insignificant. Δ means the first difference of variables.

Table III.2. Philippines Electricity Demand Model Unit root test
—Augmented Dickey Fuller Test

Sample period: 1974-2006
Critical value used in ADF test
With constant: 5 percent =-2.95; 1 percent=-3.65
With constant and trend: 5 percent=-3.55; 1 percent=-4.26
Null hypothesis: The variable has a unit root.

	Level				First difference			
	Trend	Lag length *p*	ADF statistics	Estimated γ	Trend	Lag length *p*	ADF statistics	Estimated γ
eh	Yes	0	-1.517	0.85	Yes	0	-4.611**	0.12
ep	Yes	0	-1.946	0.75	Yes	0	-4.981**	-0.33
ph	Yes	1	-2.984	0.60	Yes	0	-7.701**	-.029
pp	Yes	1	-2.370	0.75	Yes	0	-3.914**	0.29
i	Yes	0	-2.405	0.63	Yes	1	-6.341**	0.72
gdp	Yes	1	-1.749	0.90	Yes	0	-3.517*	0.40
a	Yes	2	0.693	1.03	Yes	1	-3.650**	0.34

Notes:
1 For a variable *y*, the following regression equation is considered:

$$\Delta y_t = a_1 + a_2 \cdot t + (\gamma - 1) \cdot y_{t-1} + \sum_{i=1}^{p} \beta_i \, \Delta y_{t-i} + \varepsilon_t$$

where Δ is the difference operator, t a time trend, and *p* the lag length. The ADF statistic is the t ratio on the lagged variable y_{t-1}.

2. ** indicates significant at 1 percent level and * indicates significant at 5 percent level. The null hypothesis is $\gamma = 1$ (the variable has a unit root).

Table III.3. Philippines Electricity Demand Model VAR Lag Selection 1/
(Sample period: 1974-2006)

Endogenous variables	Residential electricity consumption				Industrial electricity consumption	
	(eh, i, ph, a)		(eh, i, a)		(ep, gdp)	
	AIC	SC	AIC	SC	AIC	SC
Lag 3	**-13.522**	-10.957	**-10.877**	-9.366	-7.437	-6.704
Lag 2	-12.873	-11.041	-10.636	-9.536	-7.437	-6.888
Lag 1	-12.522	**-11.422**	-10.385	**-9.697**	**-7.458**	**-7.092**

1/ Exogenous variable are constant and debt crisis dummy variables (1984 and 1985 = 1).

Table III.4. Philippines Electricity Demand Model VAR Diagnostic Tests
(Sample period: 1974-2006)

Endogenous variables	Residential electricity consumption (eh, i, ph, a)		Residential electricity consumption (eh, i, a)		Industrial electricity consumption (ep, gdp)
Model number	I	II	III	IV	V
Selected lags	**2**	**3**	**2**	**3**	**2**
Autocorrelation LM test					
Lag order 1	21.02	30.92	14.84	12.38	0.93
Probability	*0.18*	*0.01* **	*0.10*	*0.19*	*0.92*
Lag order 2	10.80	19.98	5.63	12.11	2.18
Probability	*0.82*	*0.22*	*0.78*	*0.21*	*0.70*
Lag order 3	13.53	11.55	3.38	6.82	6.48
Probability	*0.63*	*0.77*	*0.95*	*0.66*	*0.17*
Normality	14.26	21.04	8.71	11.36	5.27
Probability	*0.08*	*0.01* *	*0.19*	*0.08*	*0.26*
Skewness	1.06	1.07	2.28	2.53	2.84
Probability	*0.90*	*0.90*	*0.52*	*0.47*	*0.24*
Kurtosis	13.21	19.97	6.43	8.83	2.44 `
Probability	*0.01* *	*0.00* **	*0.09*	*0.03* *	*0.30*
Heteroscedasticity	194.81	266.15	85.86	116.67	39.37
Probability	*0.09*	*0.23*	*0.25*	*0.41*	*0.06*

Notes:
1. Error autocorrelation is vector autocorrelation LM test for the "VAR lag" + 1th order of residual autocorrelation. Under the null hypothesis, there is no serial correlation.
2. Normality is Cholesky test. Under the null hypothesis, residuals are normal distributed.
3. Heteroscedasticity is Doornik (1996) test. Under the null hypothesis, there is no heteroskedasticity.
4. ** indicates significant at 1 percent level and * indicates significant at 5 percent level.

Table III.5. Philippines Electricity Demand Model Unrestricted cointegration analysis
(1974-2006)

Endogenous variables	Residential electricity consumption (eh, i, ph, a)		(eh, i, a)		Industrial electricity consumption (ep, gdp)
Lag order in the original VAR model	2	3	2	3	2
Model number	I	II	III	IV	V
Normalized β cointegrating vector					
eh	1.000	1.000	1.000	1.000	...
	...	...	...	...	...
ep	...	...	...	...	1.000
	...	...	...	...	...
i	-1.076	-0.931	-1.179	-0.933	...
	-3.085 ***	*-2.444* *	*-3.732* ***	*-2.743* **	...
gdp	...	...	...	...	-0.299
	...	...	...	...	*-2.435* **
ph	-0.292	-0.200	...	...	...
	-1.239	*-0.800*	...	...	...
a	-0.124	-0.191	-0.219	-0.249	...
	-1.580	*-2.096* *	*-6.426* ***	*-6.480* ***	...
constant	...	...	...	...	...
	...	...	...	...	...
time trend	...	...	...	...	-0.004
	...	...	...	...	*-3.408* **
Speed of adjustment (α) parameters					
d(eh) equation	-0.141	-0.265	-0.186	-0.299	...
	-1.560	*-2.551* *	*-1.774* *	*-2.597* *	...
d(ep) equation	...	...	...	...	-0.437
	...	...	...	...	*-2.517* **
d(di) equation	0.270	0.210	0.291	0.215	...
	4.104 ***	*3.164* ***	*3.910* ***	*2.909* **	...
d(gdp) equation	...	...	...	...	0.214
	...	...	...	...	*2.623* **
d(ph) equation	0.246	0.290	...	...	...
	2.148 **	*2.144* *	...	...	...
d(a) equation	-0.049	-0.027	-0.056	-0.028	...
	-2.019 **	*-0.902*	*-1.985* **	*-0.907*	...
Trace tests					
r=0	47.371	42.815	35.671	35.879	29.764
	[0.055] *	[0.137]	[0.009] ***	[0.009] ***	[0.016] *
r≤1	17.056	15.867	6.886	10.339	0.521
	[0.636]	[0.721]	[0.591]	[0.255]	[1.000]
r≤2	5.442	3.000	0.970	0.000	...
	[0.760]	[0.967]	[0.325]	[0.992]	...
r≤3	0.302	0.000	...	...	...
	[0.582]	[0.990]	...	...	...

Notes:
1. *Italic* indicates t statistics; [ ] indicates p value.
2. *** indicates significant at 1 percent level, ** significant at 5 percent level, and* significant at 10 percent level
3. Exogenous variable are constant and debt crisis dummy variables (1984 and 1985 = 1).

Table III.6. Philippines Electricity Demand Model Restricted cointegration analysis
(Sample period: 1974 - 2006)

Endogenous variables	Residential electricity consumption (eh, i, a)						Industrial electricity consumption (ep, gdp)
Lag order in the original VAR model	2	3	2	3	2	3	2
Model number	III	IV	III	IV	III	IV	V
LR test for parameter restrictions test	*a* is weakly exogenous		*a* is weakly exogenous, and coefficint to i =1		*a* is weakly exogenous, coefficint to i =1, and *i* is weakly exogenous,		*gdp* is weakly exogenous
Chi square	4.106	0.921	4.955	0.955	14.944	10.183	7.158
	[0.043] **	[0.337]	[0.084]	[0.620]	[0.001] ***	[0.017] **	[0.007]
Normalized β cointegrating vector							
eh	1.000	1.000	1.000	1.000	1.000	1.000	1.000
i	-1.313	-1.065	-1.000	-1.000	-1.000	-1.000	...
	-4.083 **	*-3.191* **					
gdp	...	...	...	...	...	...	-0.266
							-1.932 *
ph	...	...	...	...	...	...	...
a	-0.233	-0.257	-0.238	-0.257	-0.231	-0.222	...
	-6.705 **	*-6.818* **	*-6.895* **	*-7.023* **	*-5.343* **	*-4.854* **	
constant	...	...	...	...	...	...	...
time trend	...	...	...	...	...	...	-0.004
							-2.773 *
Speed of adjustment (α) parameters							
d(eh) equation	-0.248	-0.317	-0.268	-0.320	-0.401	-0.435	...
	-2.373 *	*-2.700* **	*-2.701* *	*-2.775* **	*-4.501* ***	*-4.362* ***	...
d(ep) equation	...	...	...	...	...	...	-0.725
	...	...	...	...	...	...	*-5.593* **
d(di) equation	0.271	0.224	0.244	0.218	0.000	0.000	...
	3.674 **	*3.009* **	*3.308* **	*2.940* **	...	...	...
d(gdp) equation	...	...	...	...	...	...	0.000
	...	...	...	...	...	...	...
d(ph) equation	...	...	...	...	...	...	...
	...	...	...	...	...	...	...
d(a) equation	0.000	0.000	0.000	0.000	0.000	0.000	...
	...	...	...	...	...	...	...

Notes:
1. *Italic* indicates t statistics; [ ] indicates p value.
2. *** indicates significant at 1 percent level, ** significant at 5 percent level, and* significant at 10 percent level
percent level.
3. Exogenous variable are constant and debt crisis dummy variables (1984 and 1985 = 1).

Table III.7. VECM: Residential Electricity Demand Model 1/ 2/
(Sample period: 1974 - 2006)

	Residential electricity consumption															
	Model R1		Model R2		Model R3		Model R4		Model R5		Model R6		**Model R7**		Model R8	
	d(eh)	d(i)	d(eh)	d(i)	d(eh)	d(i)	d(eh)	d(i)	d(eh)	d(i)	d(eh)	d(i)	d(eh)	d(i)	d(eh)	d(i)
Error correction term (-1)	-0.29	0.24	-0.29	0.23	-0.29	0.23	-0.25	0.24	-0.25	0.26	-0.26	0.23	-0.22	0.23	-0.17	0.24
	-2.44 **	*3.12* ***	*-2.48* **	*3.12* ***	*-2.69* **	*3.24* ***	*-2.43* **	*3.73* ***	*-2.53* **	*4.00* ***	*-2.76* **	*3.48* ***	*-2.22* **	*3.60* ***	*-1.89* *	*3.82* ***
deh(-1)	-0.05	-0.18	-0.05	-0.19	-0.03	-0.17	-0.05	-0.18	-0.05	-0.20	...	...	...	...	...	...
	-0.28	*-1.44*	*-0.25*	*-1.56*	*-0.16*	*-1.64*	*-0.28*	-1.72 *	-0.29	-1.85 *	...	...	...	...	...	...
deh(-2)	0.19	0.02	0.18	0.04	0.20	0.05	...	...	...	...	...	...	...	...	...	...
	0.99	*0.15*	*0.97*	*0.31*	*1.13*	*0.41*	...	...	...	...	...	...	...	...	...	...
di(-1)	-0.02	-0.46	0.00	-0.50	-0.02	-0.51	0.05	-0.50	0.04	-0.59	0.03	-0.66	0.00	-0.66	...	-0.72
	-0.08	*-2.67* **	*0.01*	*-3.09* ***	*-0.07*	*-3.37* ***	*0.20*	*-3.42* ***	*0.21*	*-4.27* ***	*0.15*	*-4.70* ***	*-0.01*	*-4.81* ***	...	*-5.75* ***
di(-2)	-0.21	-0.43	-0.22	-0.42	-0.23	-0.42	-0.17	-0.41	-0.17	-0.47	-0.16	-0.44	-0.12	-0.44	...	-0.37
	-1.08	*-3.41* ***	*-1.15*	*-3.39* ***	*-1.24*	*-3.57* ***	*-0.96*	*-3.66* ***	*-1.04*	*-4.26* ***	*-1.02*	*-3.85* ***	*-0.69*	*-3.93* ***	...	-3.56 ***
da(-1)	-0.32	0.71	-0.24	0.57	-0.27	0.55	0.00	0.62	...	...	...	...	...	...	...	...
	-0.43	*1.50*	*-0.36*	*1.36*	*-0.43*	*1.36*	*0.01*	*1.68*	...	...	...	...	...	...	...	...
da(-2)	-0.63	-0.85	-0.78	-0.60	-0.69	-0.55	-0.96	-0.61	-0.95	-0.14	-0.93	-0.03	...	...	...	...
	-0.68	*-1.41*	*-1.05*	*-1.26*	*-1.08*	*-1.33*	*-1.61*	*-1.63*	*-2.51* **	*-0.54*	*-2.55* **	*-0.12*	...	...	...	...
dph(-1)	-0.31	-0.15	-0.32	-0.17	-0.29	-0.17	-0.30	-0.17	-0.93	-0.17	-0.29	-0.14	-0.33	-0.14	-0.27	...
	-2.08 **	*-1.52*	*-0.47*	*-1.97* *	*-2.24* **	*-2.00* *	*-2.25* **	*-2.05* *	-2.25 **	-1.93 *	-2.32 **	-1.56	-2.45 **	-1.62	-2.36 **	...
dph(-2)	-0.05	0.08	...	...	...	...	...	...	...	...	...	...	...	...	...	...
	-0.26	*0.68*	...	...	...	...	...	...	...	...	...	...	...	...	...	...
Constant	-1.12	1.06	-1.10	1.04	-1.14	1.01	-0.94	1.06	0.11	1.16	-0.97	1.01	-0.92	1.01	-0.69	1.08
	-2.16 **	*3.21* ***	*-2.19* **	*3.20* ***	*-2.43* **	*3.35* ***	*-2.15* **	*3.88* ***	*3.12* ***	*4.21* ***	*-2.49* **	*3.67* ***	*-2.14* **	*3.75* ***	*-1.81* *	*3.97* ***
Debt crisis dummy 3/	0.04	-0.28	0.04	-0.28	0.04	-0.28	0.03	-0.28	0.03	-0.27	0.03	-0.28	0.06	-0.28	...	-0.29
	0.56	*-6.44* ***	*0.53*	*-6.50* ***	*0.60*	*-6.70* ***	*0.53*	*-6.85* ***	*0.55*	*-6.47* ***	*0.53*	*-6.38* ***	*0.84*	*-6.56* ***	...	*-7.20* ***
Temp	-0.01	-0.03	-0.02	-0.01	...	...	...	...	...	...	...	...	...	...	...	...
	-0.14	*-0.46*	-0.25	-0.24	...	...	...	...	...	...	...	...	...	...	...	...
Memorundam items:																
Log likelihood	115.45		114.89		114.82		113.93		111.92		109.76		105.84		103.48	

1. *Italic* indicates t statistics; [ ] indicates p value.
2. *** indicates significant at 1 percent level, ** significant at 5 percent level, and* significant at 10 percent level
3. 1984 and 1985 = 1.

Table III.8. Residential Electricity Demand Models: Sequential Reductions (F statistics)

Null hypothesis Model	Model R2	Model R3	Model R4	Model R5	Model R6	Model R7	Model R8
	F statistics						Chi square
Model R1 →	0.34	0.19	0.31	0.56	0.75	1.13	23.93
	[0.71]	[0.94]	[0.92]	[0.80]	[0.67]	[0.37]	[0.09] *
Model R2 →	---	0.04	0.31	0.66	0.88	1.33	22.83
		[0.95]	[0.87]	[0.68]	[0.53]	[0.25]	[0.06] **
Model R3 →	---	---	0.61	1.10	1.21	1.72	22.69
			[0.55]	[0.41]	[0.31]	[0.12]	[0.03] **
Model R4 →	---	---	---	1.49	1.55	2.13	20.91
				[0.24]	[0.20]	[0.07] *	[0.02] **
Model R5 →	---	---	---	---	1.67	2.43	16.88
					[0.20]	[0.06] *	[0.03] **
Model R6 →	---	---	---	---	---	3.36	12.57
						[0.05] *	[0.05] *
Model R7 →	---	---	---	---	---	---	4.71
							[0.31]

1. [ ] indicates p value.

Table III.9. VECM: Industrial Electricity Demand Model
(Sample period: 1974 - 2006)

	Production sector electricity consumption											
	Model P1		Model P2		Model P3		Model P4		Model P5		Model P6	
	d(ep)	d(gdp)	d(ep)	d(gdp)	d(ep)	d(gdp)	d(ep)	d(gdp)	d(ep)	d(gdp)	d(ep)	d(gdp)
Error correction term (-1)	-0.43	0.19	-0.42	0.20	-0.44	0.21	-0.47	0.20	-0.38	0.16	-0.36	0.19
	-2.37 **	*2.17* **	*-2.36* **	2.30 **	*-2.70* **	*2.66* ***	*-3.03* ***	*2.66* **	*-2.79* ***	*2.48* **	*-2.70* **	*2.99* ***
dep(-1)	-0.09	-0.05	-0.07	-0.04	-0.07	-0.04	...	...	...	...	...	...
	-0.54	*-0.57*	*-0.45*	*-0.53*	-0.47	-0.52	...	...	...	...	...	...
dgdp(-1)	0.30	-0.13	0.35	-0.12	0.35	-0.12	0.33	-0.14	...	...	...	...
	0.99	*-0.91*	*1.20*	*-0.89*	*1.23*	*-0.91*	1.19	-1.02	...	...	...	...
ppgap	*0.11*	-0.05	*-0.13*	-0.06	*-0.14*	-0.05	*-0.13*	-0.05	*-0.15*	-0.04	*-0.12*	...
	-1.21	-1.14	*-1.86* *	-1.63	*-2.02* *	-1.58	*-2.00* *	-1.51	*-2.39* ***	-1.28	*-2.71* **	...
ppgap(-1)	*-0.03*	-0.01	...	...	...	...	...	...	...	...	...	...
	-0.47	-0.26	...	...	...	...	...	...	...	...	...	...
Constant	1.68	-0.73	1.62	-0.75	1.72	-0.80	1.81	*-0.75*	1.46	*-0.61*	1.38	*-0.71*
	2.37 **	*-2.15* **	*2.36* **	*-2.27* **	*2.70* **	*-2.63* ***	*3.04* ***	*-2.63* **	*2.79* ***	*-2.44*	*2.71* **	*-2.94* ***
Debt crisis dummy	0.009	-0.13	0.01	-0.12	0.01	-0.12	0.01	-0.13	-0.004	-0.12	...	*-0.13*
	0.17	*-5.22* ***	*0.25*	*-5.34* ***	*0.27*	*-5.43* ***	*0.22*	*-5.59* ***	*0.93*	*-5.51* ***	...	*-8.37* ***
Temp1	0.02	-0.01	0.02	-0.01	...	...	...	...	...	...	...	...
	0.42	*-0.52*	0.44	-0.53	...	...	...	...	...	...	...	...
Memorundam items:												
Log likelihood	142.05		141.90		141.11		140.93		137.52		136.47	

Notes:
1. *Italic* indicates t statistics; [ ] indicates p value.
2. *** indicates significant at 1 percent level, ** significant at 5 percent, and * significant 10 percent level.
3. 1984 and 1985 = 1.

Table III.10. Industrial Electricity Demand Models: Sequential Reductions (F statistics)

Null hypothesis Model	Alternative model					
	Model P1	Model P2	Model P3	Model P4	Model P5	Model P6
	F statistics					Chi-square statistics
Model P1	---	0.10 [0.90]	0.35 [0.84]	0.28 [0.94]	0.88 [0.53]	11.17 [0.34]
Model P2	---	---	0.62 [0.54]	0.38 [0.82]	1.18 [0.32]	10.88 [0.21]
Model P3	---	---	---	0.15 [0.86]	1.49 [0.21]	9.30 [0.16]
Model P4	---	---	---	---	3.10 [0.06] *	8.92 [0.06] *
Model P5	---	---	---	---	---	2.11 [0.35]

1. [ ] indicates p value.

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