Money Demand and the Welfare Cost of Inflation:

Empirical Evidence from East Asian Countries

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Abstract

This dissertation presents empirical evidence of the long-term relationship between money (M1) and inflation from five selected East Asian countries during the period from 1977–2014. The two money demand specifications, namely, the semi-log and the log-log functional forms initiated by Cagan (1956) and Meltzer (1963), respectively, are examined using time-varying and structural break tests for stationary and cointegration before estimating cointegration equations with ordinary least squares (OLS) regression. In addition to estimating the money demand function, the welfare cost of inflation is calculated following Bailey's (1956) consumer surplus approach. Results indicate that there is an insignificant effect of inflation in the five countries, as well as in comparison to existing studies. On the other hand, there is a distinction between the two specifications, with the log-log form often producing more consistent and reasonable results.

Keywords: The Welfare Cost of Inflation, Money Demand, Structural Break test, Asia

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Attestation of authorship

I hereby declare that this submission is my own work and that, to the best of my knowledge and belief, it contains no material previously published or written by another person (except where explicitly defined in the acknowledgements), nor material which to a substantial extent had been submitted for the award of any other degree or diploma of a university or other institution of higher learning,

Signed ________________________________ Date_______25/07/16_________________

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I. Introduction

Inflation is a prevailing issue in the contemporary macroeconomic literature. The cost of inflation to social welfare has been investigated by various studies across different countries and regions in the world. This welfare cost is defined as “a tax on the holding of cash balances, a cost which is fully analogous to the welfare cost (or excess burden) of an excise tax on a commodity or productive service” (Bailey, 1956). Bailey (1956) initiates the method for estimating the welfare cost of inflation, which links inflation and real money balances together and treats real money balances as a consumption good in a fluctuating interest rates model. This so-called, the “consumer surplus approach”, measures the welfare cost as the area under the money demand curve as the interest rate rises from r_0 to r_1 (see figure 1).

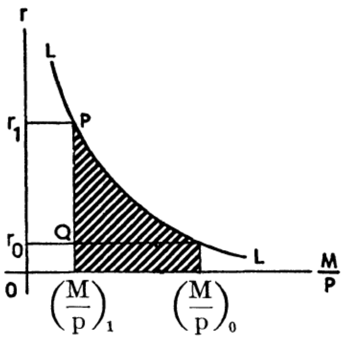

Figure 1: Average welfare cost

A preliminary step to computing welfare cost is to estimate a money demand function that determines the real relationship between money balances and inflation rates. There is a diverse spectrum of money demand models depicting this relationship. On the one hand, the concern is the selection of appropriate variables. Besides sharing common important variables such as money aggregates, income or interest rates, they bring forth the relationship between the quantity of money demanded and a set of a few important economic variables linking money to the real sector of the economy (Judd and Scalding, 1982). Sriram (2001) generalises the functional long-term relationship as:

$$\frac{M}{P} = f(S, OC) \qquad (1.1)$$

where the demand for real balances (M/P) is a function of the chosen scale variable (S) to represent the economic activity and the opportunity cost of holding money (OC). M stands for the selected monetary aggregate in nominal terms and P for the price. The scale variable is used in the estimation as a measure of transactions relating to the economic activity. It is usually represented by variables expressing income, expenditure, or wealth concepts, while the opportunity cost variables may involve expected inflation rates, domestic and foreign interest rates. Specifically to this dissertation, real gross domestic product (GDP) is chosen as the scale variable, money aggregate as M and nominal deposit interest rates as the opportunity cost variable. However, in order to focus on the relationship between money demand and interest rate, M is divided by GDP, representing the fraction of money balances to income.

On the other hand, the selection of a proper framework is also subject to variation across studies over time. The long-term relationship of money demand can be described in either level form or logarithmic form. "The economic theory does not provide any rationale as to the correct mathematical form of the money demand function. There is consensus, however, that the log-linear version is the most appropriate functional form." (Zarembka, 1968) Despite this theoretical assumption, it is contemporaneously lacking of empirical evidence for further consideration. Recognising this fact, this dissertation provides an empirical comparison between the semi-log and log-log money demand specifications to strengthen the conclusion.

Regarding the estimation technique, the traditional partial adjustment method has been replaced by the error-correction approach. The error-correction model (ECM) is extremely useful and common in dealing with univariate as well as multivariate time series functions. By pre-testing individual time series with unit root testing techniques such as Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) or Zivot and Andrews

(ZA), the ECMs have solved the autocorrelation problem caused by the deterministic stationary process.

Furthermore, a structural break test can also be incorporated in the pre-estimation process to reduce spurious correlations that may occur due to a regime shift in the time series function. A structural break test technique such as Gregory and Hansen (GH, 1996) can help to identify an unknown break point that causes a one-time change in the intercept or the slope (or both) of the money demand function. After examining the stationarity and stability of the model, the correct model specification is rendered and estimated for the coefficient using regression techniques.

The main contribution of this dissertation is to provide evidence of the welfare cost of inflation in Asian regions using Bailey's (1956) method. The study sheds additional light on this issue by incorporating a structural break technique in the computation process. To my best knowledge, there have not yet been any studies estimating the welfare cost based on structural break analysis.

This dissertation has the following structure. The next section gives a brief review of the literature including theoretical development and empirical evidence of the welfare cost of inflation. Section 3 provides details of the data and method implemented by the dissertation. The empirical results are presented and discussed in Section 4. Finally, Section 5 offers conclusions.

II. Literature review

1. Theoretical framework

1.1 Money demand in a general framework

In general, demand for money is demand for real balances. The three motivations for the demand for money are transaction, precautionary and speculative. Early economists have focused solely on the transaction motive, which led to the establishment of "transactions theories" on money demand. Fisher (1911) provides the famous equation, the so-called "equation of exchange", $M_S V_T = P_T T$, which relates the quantity of money in circulation M_S to the volume of transactions T and the price level of articles traded P_T in a given period through a proportionality factor V_T called the "transactions velocity of circulation".

The alternative approach to the quantity theory of money, the so-called "Cambridge approach", was primarily developed by Pigou (1917) and Marshall (1923). The formula is written as: $M_d = kPy$, where the demand for money terms (M_d) is proportional to the nominal level of income (Py) for each individual and for the aggregate economy as a whole. It is recognised that k might depend on other variables in the consumer allocation problem, such as the interest rates and wealth. (Sriram, Survey of Literature on Demand for Money: Theoretical and Empirical Work with Special Reference to Error-Correction Models, 1999c) Although the speculative motive is mentioned in the factor k, the main focus is solely on transactions.

The modern theory of money demand is illuminated by the Keynes theory, which inclusively considers other functions of money. Keynes (1936) states the function as: $M_d = f(y, i)$, where the demand for money balances is a function of income (y) and interest rates (i). This money demand formulation has become the primary tool for researchers in analysing the demand for money up to the present.

As opposed to the Keynesian view, Friedman (1956) takes for granted the motive for holding money as an assumption that money is a durable good which yields a flow of unobservable services. The canonical equation can be written as: $\frac{M^d}{P} = \Phi(y_p)$, which indicates that the quantity of money is only determined by the real permanent income. Friedman's theory implies that the money demand function is highly stable and insensitive to interest rates.

Cagan (1956) introduces the model of adaptive expectation using the expected rate of inflation in a semi-log form in an analysis of the hyperinflation episode in several European countries from 1929–1952. The canonical specification of the money demand function can be written as:

$$m_t = c_t - \alpha r_t - \alpha \pi_t^e \quad (2.1)$$

where m_t = the log of real money, c_t = real consumption, r_t = the real interest rate and π_t^e = expected rate of inflation. The concept of rational expectation is later enhanced by Lucas (1972, 1973, 1976), Sargent and Wallace (1975), Barro (1976), and Barro and Gordon (1983).

Nevertheless, Cagan's model experiences some drawbacks, as Lucas (1976) pointed out in his paper. Lucas's critique rests on the fact that the model is built on public expectation of inflation, which is always subject to change, and that we should avoid using equations that will tend to shift with policy changes. Lucas suggested to improve the conventional technique for policy evaluation by constructing a model based on structural parameters, which are invariant with respect to policy intervention.

Beside the rational expectation model, the OLS model offers an alternative method to estimate the money demand function. The merit of this approach is that its result is formulated based on testable hypotheses in a cointegration framework. The two most popular specifications that have been used in the empirical estimation of money demand functions are the log-linear and double-log functions of the form. In particular, the log-linear money demand function is often expressed as:

$$\ln\left(\frac{M}{P}\right) = \alpha + \beta lnY + \gamma R \qquad (2.2)$$

This log-linear form is criticised by Serletis (2007) as "simply unbelievable in the volatile financial environment in which we find ourselves" and "inappropriate for monetary policy purposes". Researchers, hence, favour the double-log form in empirical study in an equation of the form:

$$\ln\left(\frac{M}{P}\right) = \alpha + \beta lnY + \gamma lnR \qquad (2.3)$$

Many of the early studies tended to ignore the time series property of the variables and treated them as time-invariant variables. This dynamic aspect of the money demand specification is later posited by the partial adjustment model. This model posits the existence of a desired level of real money balances, reflecting what real money demand would be if there were no adjustment costs, and further assumes that the actual level of money balances adjusts in each period only part of the way toward its desired level. (Serletis A. , 2007) A reflection of the short-run demand for money function can be obtained [e.g., Chow (1966) and Goldfeld and Quant (1973)] as:

$$\log\left(\frac{M_t}{P_t}\right) = \lambda\alpha + \lambda\beta_1 logY_t + \lambda\beta_2 R_t + (1-\lambda)\ log\left(\frac{M_{t-1}}{P_{t-1}}\right) + e_t \qquad (2.4)$$

where (M_t/P_t) denotes the actual value of real money balances, e_t is a random error term, and λ is a measure of the speed of adjustment, with $0 \leq \lambda \leq 1$; $\lambda = 1$ corresponds to full immediate adjustment while smaller values represent slower, more sluggish, adjustment. The shortcoming of the partial adjustment framework is that it is a seriously misspecified and highly restrictive dynamic (see Cooley and LeRoy, 1981; Goodfriend, 1985; Hendry, 1979 and 1985; Hendry and Mizon, 1978). To counter this problem, ECM can be used to test for the stationarity of the variables before selecting the appropriate estimation techniques.

The ECM is shown to contain information on both the short- and long-run properties of the model with disequilibrium as a process of adjustment to the long-run equilibrium. (Sriram, 2001) The early ECM on money demand was developed by Engle and Granger (1987) with a single equation cointegration relationship between money and scale variables. Johansen (1988) and Johansen and Juselius (1990) later generalised the ECM to allow for multiple cointegration vectors between non-stationary variables in order to provide a fuller characterisation of the long-run determinant of demand.

There is a vast literature on the approach to the stationary (unit root testing) issue [see Phillips–Perron (1988), Zivot and Andrews (1992) and Stock (1994)]. One of the very popular unit root tests was introduced by Dickey and Fuller (1979), in which a variable is tested for stationarity using OLS to estimate the coefficients of the model:

$$\Delta y_t = \alpha_1 y_{t-1} + e_t \tag{2.5}$$

where Δy_t is the first difference, and e_t is an independently and identically distributed zero-mean error term. The null hypothesis of $\alpha_1 = 0$ is examined using a standard t-test in OLS to determine whether y_t is a pure random walk and thus nonstationary. The model is further generalised in an augmented Dickey-Fuller test to control for serial correlation where:

$$\Delta y_t = \sum_{j=1}^{p} a_j y_{t-j} + e_t \tag{2.6}$$

Another issue with the time series function is the breaking trend of the linear relationship due to a one-time change in the intercept or in the slope (or both) of the trend function. Perron (1989) argues that certain ‘big shocks’ do not represent a realisation of the underlying data generation mechanism of the series under consideration and that the null should be tested against the trend-stationary alternative by allowing, under both the null and the alternative hypotheses, for the presence of a one-time break (at a known point in time) in the intercept or in the slope (or both) of the trend function. Indeed, Gregory, Nason and Watt (1994) have shown that the power of conventional ADF tests can fall sharply due to the presence of a structural break.

In particular, Perron (1989) modifies the ADF regression by adding dummy variables DU_t, DT_t and $D(T_B)_t$ to the model:

$$y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_B)_t + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + e_t \qquad (2.7)$$

where $DU_t = 1$ and $DT_t = t$ if $t > T_B$ and 0 otherwise, and $D(T_B)_t = 1$ if $t = T_B + 1$ and 0 otherwise. T_B (with $1 < T_B < T$, where T is the sample size) denotes the time at which the change in the trend function occurs. Perron's (1989) assumption on the endogenous break point has been criticised by Christiano (1992) as it may lead to misspecification, thus the break point should be correlated with the data.

Gregory and Hansen (1996) extend the issue of structural break by constructing a test procedure that does not require information regarding the timing of a break. The test's null hypothesis assumes no cointegration against the alternative of cointegration with a single shift at an unknown point in time. Further illustration of this method will be presented in section 3 of the dissertation.

1.2 The use of money demand in welfare estimation

In the context of the welfare cost of inflation, money demand estimation is a crucial step prior to calculating the welfare function. There are several specifications of money demand functions that have been utilised by researchers applying different estimation techniques. Bailey (1956) and Friedman (1969) use the semi-log specification replicated from Cagan's (1956) equation (2.1) with a money demand function of the form:

$$\frac{M}{P} = e^{\propto E + \gamma} \qquad (2.8)$$

where E is the expected rate of inflation and where α and γ are constants. This specification implies a monotonically decreasing impact of inflation on the demand for real cash balance as the rate approaches zero. Lucas (2000) modifies Cagan's (1956) as well as Meltzer's (1963) money demand specifications as:

$$\ln(m) = \ln(B) - \varepsilon r \qquad (2.9)$$

$$\ln(m) = \ln(A) - \eta \ln(r) \qquad (2.10)$$

where B > 0 and A > 0 are the intercepts, ε > 0 and η > 0 are the short-term interest rates, semi-elasticity and elasticity of money demand, respectively.

1.3 Welfare cost estimation

The consumer surplus approach defines the welfare cost of a nominal interest (r) as the income compensation needed to leave consumers indifferent to the choice between living in the steady state of a constant interest rate r and otherwise living in the steady state of a zero interest rate. Bailey (1956) associated the welfare cost with the changes in real cash balances and computed this cost by integrating the area under the money demand curve as the interest rate rises from 0 to r > 0 to find the lost consumer surplus, then subtracting off the seigniorage revenue r_m to isolate the deadweight loss. His equation is shown below:

$$W(r) = \int_{k_0^*}^{k_E^*} r(k)dk \qquad (2.11)$$

where k_0^* and k_E^* are defined as the steady state and the inflation state of the interest rate, respectively. Lucas (2000) shows that when the money demand take the log-log form, the welfare cost of inflation is expressed as:

$$W(r) = A\left(\frac{\eta}{1-\eta}\right) r^{1-\eta} \qquad (2.12)$$

where A and η are the exponential of the intercept and the elasticity of interest rates respectively and taken directly from equation (2.10). When the money demand takes the semi-log form, then the welfare cost is:

$$W(r) = \frac{B}{\varepsilon}[1 - (1 + \varepsilon r)e^{-\epsilon r} \qquad (2.13)$$

where B and ε are the slopes of the intercept and interest rates taken directly from equation (2.9). Lucas (2000) also proposed an alternative method, so-called the “compensating variation approach”, for estimating this welfare cost using the Sidrauski (1967) framework. In particular, the welfare cost of the nominal interest rate R, w(R), which is defined as the percentage of income compensation needed to leave the household indifferent between positive and zero nominal interest rates, is expressed as:

$$w(R) = \exp\left[-\frac{\eta \ln\left(\frac{1}{A(Rexp(\eta lnR)) - \frac{\eta}{A(\eta+1)} - \frac{1}{A(\eta+1)}}\right)}{\eta+1}\right] - 1 \qquad (2.14)$$

where A and η are described as above in equation (2.13). Another alternative framework to estimate the cost of inflation is proposed by Burstein and Hellwig (2008), focusing on the choice-making behaviour at the microeconomic level. As consequences of exogenous fluctuations in price levels, household and firms adjust their consumption and input respectively, which subsequently affects consumer and producer surplus. The variation of the social welfare, measured as the sum of consumer utility and producer profit, represent the welfare loss/gain caused by inflation. Burstein and Hellwig’s utility function and production function are expressed as:

$$U = log\left(\left[bC^{(\eta-1)/\eta} + (1-b)\left(\frac{M}{P}\right)^{(\eta-1)/\eta}\right]^{\eta/(\eta-1)}\right) + \psi\log(1-N) \qquad (2.15)$$

where C, $\frac{M}{P}$, N denote consumption goods, real money balances and labour supply respectively.

$$\pi_{it} = p_{it}c_{it} - W_t\left(\frac{c_{it}}{z_{it}}\right)^{1/\alpha} \qquad (2.16)$$

where π = a firm’s nominal profit, c = consumption, p = price, W = wages, and z = idiosyncratic cost shock to a firm’s productivity. By altering the assumptions of the welfare estimates using different general equilibrium frameworks, Lucas (2000) and Burstein and Hellwig (2008) have contributed to the understanding of the cost of inflation and inspired further investigation in the field.

2. Empirical evidence

2.1 Money demand

2.1.1 Choices of variables

One of the most important aspects of modelling the money demand function is the selection of appropriate variables, including money stock variables, scale variables and opportunity cost variables. Various empirical studies differ due to their theoretical base, which can impact the results and outcomes greatly.

Money stock variable

The most common variables representing the stock of money are M0 and M1, which are classified as narrow money, while broader money variables such as M2, M3 and M4 are arguably less favourable as they might muddy the interest rate effect. In particular, Hossain (1994) compares the performance of the narrow money demand function and the broad money demand function, in which the empirical results indicate that narrow money is a more stable and more appropriate instrument for policy makers.

However, as a result of the evolving financial system with the extended boundaries of narrow money, broad money is considered as becoming more stable and appropriate for use in evaluating the long-run economic impact of changes on monetary policy. Hafer and Jansen (1991) find that M2 is more stable, hence, it is the preferable measure over M1 aggregate.

Besides using individual monetary aggregate, the method in which the monetary asset is aggregated based on equal distribution of monetary components has been utilised in the form of:

$$M_t = \sum_{j=1}^{n} x_{jt} \tag{2.17}$$

where x_{jt} is one of the n components of the monetary aggregate M_t. This summation index implies that all monetary components contribute equally to the money total, and it views all components as dollar for dollar perfect substitutes. (Serletis A. ,

2007) This method, however, is biased toward the validity of the structural economic variable for the services of the quantity of money.

More recently, studies have been using the so-called Divisia aggregate, which was derived by Barnett (1980), who links economic monetary policy, aggregation and index number theory together in an attempt to provide properly weighting monetary components. The Divisia index is defined as:

$$logM_t^D - logM_{t-1}^D = \sum_{j=1}^{n} w_{jt}^*(logx_{jt} - logx_{j,t-1}) \qquad (2.18)$$

where the growth rate of the aggregate money demand is the weighted average of the growth rates of the components, with the Divisia index (w_{jt}) being the expenditure share of n assets during period t.

Scale variable

There is a deviation in the choices of the scale variable based on the two theories on the motive for holding money, namely, transactions theories and asset theories. While the transactions theories of money demand focus on the level of income as the relevant scale variable, the asset theories place more emphasis on wealth. In the empirical estimation, however, wealth is a complicated concept which is often difficult to measure.

The level of income, on the other hand, is more commonly used to represent the scale variable. Cagan's (1956) model of adaptive expectation, as discussed previously, can be applied to measure expected income. The shortcoming of this approach falls on the rationality, which only implicitly analyses information provided by the model, used by economic agents in formulating expectations of the future. The method has been embraced by number of researchers such as Lucas (1972, 1973), Sargent and Wallace (1975), and Barro (1976).

Alternative variables for the level of income such as GNP (gross national product), GDP and GNI (gross national income) have been heavily used in the contemporary empirical studies. Judd and Scadding (1982) criticises the use of GNP to represent the level of income as being exclusive on transactions in financial assets, sales of intermediate goods, transfers, and purchases of existing goods, all of which are likely to affect the demand for money. Whereas Bomberger and Makinen (1980) recommend GNI, which is GNP plus the term of trade for an open economy, as to reflect the impact of foreign trade on domestic transactions. However, the majority of studies utilise GDP as a relevant scale variable due to availability of the data.

Opportunity cost variables

The opportunity cost variables imply the difference in the rate of return between holding cash balances and alternative assets. The former, which is the own rate of money, is often assumed to be zero, as otherwise it may be difficult to obtain the true rate. The latter involves diverse instruments representing different type of alternative assets or monetary theories. For example, according to the transaction theories one may use short-term interest rates such as Treasury bills, commercial paper or saving deposit rates; whereas those that adapt the asset approach may apply long-term rates.

Besides the variables that have been discussed, there are other variables that may play a role in the money demand function. Kumar, Webber and Fargher (2010), for instance, use the exchange rate and inflation rate beside the short-term interest rate as proxies for the cost of holding money. The specification utilises the use of the semi-log function of the form:

$$\ln(m_t) = \theta_0 + \theta_y \ln(y_t) + \theta_R R_t + \theta_E \ln(E_t) + \theta_\pi \pi_t \tag{2.19}$$

where m = real narrow money stock, y = real output, R = cost of holding money proxied by the nominal short-term interest rate, E = cost of holding money proxied by the real effective exchange rate, π = cost of holding money proxied by the inflation rate.

There is a new trend in money demand modelling strategies toward using long-term interest rates instead of short-term rates in tracking the opportunity cost of holding money [see, for example, Anderson and Rasche (2001a, b), Dutkowsky and Atesoglu (2001), and Dutkowsky and Cynamon (2003)]. This can be justified if the expectation of future short-term interest rates has a larger effect on money demand than the real short-term rates. However, there is an internal inconsistency with using long-term interest rates when Divisia aggregates or other disaggregated money is used for the money stock variable [see Mahoney (1988), Mehra (1991), Anderson and Collins (1998), Dutkowsky and Dunsky (1998)].

2.1.2 Cross-country evidence

In general, the elasticity of interest rates obtained by empirical tests is often small in magnitude and negative in sign, which conforms to the Friedman (1969) rule that the demand for money function is highly stable and insensitive to interest rates, regardless of the estimation techniques. For instance, Darrat (1986) estimated the money demand equation for three OPEC members: Saudi Arabia, Libya and Nigeria, during the period 1963–1970 using an OLS method that provides results for interest elasticity between -0.002 and -0.01. This estimation is consistent with other studies using different techniques, such as Fielding (1994), Nwaobi (2002) and Akinlo (2006), which apply Johansen maximum likelihood (JML), vector autoregression (VAR) and autoregressive distributed lag (ARDL) respectively.

There seems to be a deviation in the magnitude of interest rates elasticity in a comparison between developing and developed countries. Developing countries with a high level of inflation often yield interest elasticity (in absolute value) of between 0.01 to 0.0001, while developed countries with a stable economy tend to have a higher interest elasticity (in absolute value) of between 0.1 to 0.9. For example, Serletis and Yavari (2007) estimate the elasticity of interest rates for seven developed European countries from 1960 to 2000 with a value of -0.22 for Belgium, -0.55 for Austria, -0.10 for France, -0.24 for Germany, -0.35 for the Netherlands, -0.52 for Ireland and -0.23 for

Italy. In contrast, Salha and Jaidi (2014) estimate this parameter for Tunisia – a developing country – from 1979 to 2011 as -0.07. These findings imply that the higher the level of inflation, the more insensitive the money demand function is to interest rates.

In another aspect, there is a lack of comparison between the log-log and semi-log specifications on the magnitude of interest elasticity. Lucas (2000) endogenously determined the interest elasticity for the two specifications. For the log-log specification, he examined the money demand function at 0.3, 0.5 and 0.7 as the values for interest elasticity, and for the semi-log specification, the interest elasticity values are 3, 5 and 7. Lucas (2000) found the values of 0.5 for the log-log function and 7 for the semi-log are best fit to the data. Ireland (2009) is one of the few studies to examine both specifications using the regression method. The interest elasticity is estimated at 0.089 for the log-log form and at 1.79 for the semi-log. These results confirm Lucas's (2000) assumption made on the values of interest elasticity.

2.2 Welfare estimation

2.2.1 The US

Most of the earliest works on the welfare cost of inflation pointed to the evaluation of the Federal Reserve policy on keeping a low but positive inflation in the US in the 20th century. Dotsey and Ireland (1994) estimated that a 4% sustained inflation would cost the economy an equivalent of .41% output per year. These estimates indicated an appropriate price stability path should be made a principal objective for monetary policy. Further, the findings demonstrated the usefulness of the general equilibrium model for policy evaluation. A partial equilibrium approach to measuring the welfare cost of suboptimal policy was proved to underestimate the welfare effects.

Eckstein and Leiderman (1991) showed that an annual rate of 10% inflation results in a loss of utility equivalent to about 1% in GNP. This is more than double the .39% of GNP figure computed by Cooley and Hansen (1989) for the US and the .3% and .45% of GNP figures reported by Fischer (1981) and Lucas (1981) respectively. In addition, the estimated model suggested a relationship between the welfare cost and the degree of risk aversion. With other things being equal, the higher the degree of risk aversion, the lower the welfare cost.

Ireland (2009) replicated Meltzer's specification of the money demand function with a log-log form and estimated the welfare cost of a 10% change in inflation from .03% to .13%. The dissertation also estimated the welfare cost based on Bailey's semi-log money demand function. Interestingly, the findings provided estimates of the cost of a 10% inflation that lie between 0.20 and 0.22% of income; these numbers are still smaller than, but resemble more closely, Fischer's (1981) and Lucas's (1981) estimates.

Silva (2012) used bond market and bond trading frequency to estimate the welfare cost of inflation. According to the paper, when the timing of bond market trades is endogenising, the elasticity of money demand increases and causes an increase in the welfare cost of inflation. Further, an estimation of the welfare cost was constructed in terms of nominal GDP by fluctuating the inflation rate from 0 to 10%. The result showed an approximation cost of $US100 billion, which corresponds to 1% of welfare costs.

Miller et al. (2014) reconsidered the welfare cost of inflation for the US economy using time-varying variables and testing for unit root instead of the traditional time-invariant method. The results estimated a welfare cost in the range of 0.025% to 0.75% of GDP and averaging 0.27% for a 10% inflation rate. These estimates fall in the ranges of existing studies reviewed above.

On the other hand, the reduction of inflation can resemble a welfare gain for the economy. Martin Feldstein (1996) estimated a perpetual welfare gain equal of about 1% of GDP a year by shifting the equilibrium rate of inflation from 2% to 0. Miquel Faig and Zhe Li (2006) found that the welfare gains of eliminating the United States' monetary business cycle observed from 1892 to 2005 is 0.01% of GDP, while the welfare gains of reducing the observed average rate of inflation to the Friedman (1969) rule is 0.26% of GDP.

2.2.2 European countries

Kimbrough and Spyridopoulos (2012) measured the welfare cost of inflation for the case of Greece, using quarterly data from 1980 to 1990. Log-log and semi-log money demand functions were estimated using both OLS and Stock and Watson's (1993) dynamic OLS method. These estimates showed the welfare cost of a 10% inflation rate in the range of 0.59% to 0.91% of GDP. The finding is significantly higher than the welfare cost found in the US and developed countries.

Serletis and Yavari (2005) estimated the welfare cost of inflation in Italy using an annual database for the period from 1861 to 1996. The findings showed that reducing the interest rate from 14% to 3% would yield a benefit equivalent to an increase in real income of 0.4%, a relatively small welfare gain.

Later work by Serletis and Yavari (2007) extended the research to seven European Union (EU) countries - Austria, Belgium, France, Germany, the Netherlands, Ireland, and Italy - using data for the pre-EU period and an approach inspired by Bailey (1956) and Lucas (2000). In a comparison of welfare cost estimates across the different countries, a 5% reduction in interest rates (from 10 to 5%) would yield a benefit equivalent to an increase in real income of about 0.3% in Belgium, 0.45% in Austria, 0.1% in France, 0.2% in Germany, 0.4% in the Netherlands, 0.5% in Ireland and 0.004

(0.4%) in Italy. These results indicated that in big countries, like France and Germany, the welfare cost of inflation is much lower than in small countries.

Dibooglu and Kenc (2008) examined the welfare costs of inflation in a stochastic general equilibrium balanced growth model. A welfare gain in the magnitude of 21.16% of initial capital was captured by a monetary policy that brings down inflation from the baseline (3.4%) to the optimal rate (-9.54%). While there is no real data that supports the benefits of inflation reduction, it is strongly supported by theoretical models showing that a decline in level of inflation would save consumers, as well as the economy, a considerable amount of money.

2.2.3 African countries

Gupta and Uwilingiye (2008) applied an alternative approach for estimating money demand, using the long-run horizon regression proposed by Fisher and Seater (1993) to estimate the long-run equilibrium relationship between money balance as a ratio of income and the Treasury bill rate for South Africa over the period from 1965 to 2007. Results indicated a significant difference from the welfare cost estimated by Bailey’s and Lucas’s model, especially at a higher rate of inflation. The results showed a much higher figure for welfare cost compared to other developed countries, around 1.08% of GDP at 10% inflation.

Makochekanwa (2008) estimated the long-run equilibrium relationship between money balance as a ratio of income and the Treasury bill rate for the Zimbabwean economy during the period from 1980 to 2008, using both quarterly and monthly data. The paper incorporated a unit root test and a Johansen test for stability. Estimates suggested that the welfare cost of inflation for Zimbabwe ranges between 0.9% and 23.4% of GDP for a band of 10 to 300% inflation for the entire period and 0.4% and 27.6% for the hyper-inflation period.

2.2.4 Alternative frameworks

Apart from the similar method reviewed above, Burstein and Hellwig (2008) applied new quantitative models for quantifying the benefits of low inflation. The results estimated a welfare cost of 2.2% consumption for a steady-state annual inflation rate of 10%. Further, the effect of menu cost was analysed to be negligible with a welfare cost of less than 0.1% relative to the opportunity cost of real money balances. In comparison between an economy with Calvo-style price staggering and a menu cost economy, Calvo-style staggered pricing raises the welfare costs of inflation by anywhere from 0.1% to 5% relative to the menu cost model.

Pablo Guerrón-Quintana (2008) applied a different approach using a micro-founded macro-econometric model, which found that a 10% inflation entails a steady state welfare cost of 2% of annual consumption. This result is slightly larger than that reported by Lucas (2000) due to the presence of frictions like habit formation that tend to amplify the welfare cost. Taking into account transitional effects, the cost drops to 1.2%. Under some circumstances, the transitional effects can erase most of the steady state welfare loss. On the other hand, the total welfare gain was estimated at 1.2% of annual consumption for a decline of 10 percentage points in inflation. This result suggests that environments with low inflation are desirable. The result is even more appealing as it comes from the type of models now being used for policy analysis around the world.

In the context of labour market and welfare, Arbex and O'Dea (2014) showed different welfare costs of inflation under different labour market structures. In particular, the welfare cost is larger in the presence of social networks with a welfare cost of .2132% of the economy's output for a 10% inflation compared to .1878% in an empty network economy. In comparison with other labour market structures from different studies, for instance, in the Walrasian labour market of Cooley and Hansen

(1989), the welfare costs of a reduction from an annual inflation rate of 10% to the optimal rate is approximately 0.1 - 0.4 % of GNP. Assuming labour market frictions in the form of search unemployment, Heer (2003) shows that a reduction of benchmark money growth of 1.3% to 0 results only in a small utility gain of .03%.

III. Data and method

1. Data

The empirical work of this dissertation utilises annual data for GDP (as constant local currency unit, LCU), deposit interest rates and money aggregate (as current LCU) over the period from 1977 to 2014 for five selected Asian countries: Indonesia, Japan, Korea, the Philippines and Thailand. Raw data for the five countries were constructed using World Development Indicators from the World Bank database. This sample period is constrained by the availability of data, which is restricted to only annual data. Therefore, there only 38 observations for each country and total of 190 observations were acquired for the study.

Given the fast-growing diversity of other types of money such as M2 and M3, the main focus of this dissertation is the investigation of the relationship between the demand for real-cash balances and interest rates, hence, M1 is chosen as the money aggregate variable, representing the most narrowed form of money. On the other hand, the selection of the five specified countries above is partly due to the availability issue. However, it is also intentional as these are the most stable economies in the region in which the market is freely behaved.

Descriptive statistics for the five countries are presented in table 1 below. The pre-model data shows a big gap in the level of GDP, money aggregates and interest rates between the five countries. In particular, Indonesia is the biggest economy with the mean of GDP equal to 3939.61 trillion in LCU; following are Korean and Japan with the mean of GDP equal to 687.23 and 423.53 trillion in LCU, repectively; while Philippines and Thailand are much smaller with the mean of GDP only equal to 3.57 and 2.7 trillion in LCU, respectively. However, money aggregates in the five countries are not proportional with the size of the economy. The average money aggregates of Japan and Indonesia are biggest, equal to 274.27 and 246.46 trillion in LCU, repectively; following are Korean, Thailand and Philippines with the average money aggregates of 146.19, 0.53 and 0.49, respectively. The volatility of interest rates is also varied greatly among these countries with Japan is the least and Indonesia is the most volatile. The standard deviation of interest rates for Japan, Thailand, Korea, Philippines and Indonesia are 1.6, 4.36, 4.41, 5.07 and 6.97, respectively.

Table 1: Descriptive statistic

Variables	Indonesia	Japan	Korea	Philippines	Thailand
GDP (in trillion LCU)	μ = 3939.61 σ = 2059.9 $\tilde{x}$ = 3913.03	μ = 423.52 σ = 90.94 $\tilde{x}$ = 459.89	μ = 687.23 σ = 405.2 $\tilde{x}$ = 659.19	μ = 3.57 σ = 1.43 $\tilde{x}$ = 3.09	μ = 2.7 σ = 1.36 $\tilde{x}$ = 2.81
Money aggregates (in trillion LCU)	μ = 246.46 σ = 341.4 $\tilde{x}$ = 52.05	μ = 274.27 σ = 182.11 $\tilde{x}$ = 199.77	μ = 146.19 σ = 179.13 $\tilde{x}$ = 35.31	μ = 0.49 σ = 0.61 $\tilde{x}$ = 0.21	μ = 0.53 σ = 0.48 $\tilde{x}$ = 0.37
Interest rates (%)	μ = 13.01 σ = 6.97 $\tilde{x}$ = 11.96	μ = 1.71 σ = 1.6 $\tilde{x}$ = 0.85	μ = 8.54 σ = 4.41 $\tilde{x}$ = 8.25	μ = 9.28 σ = 5.07 $\tilde{x}$ = 8.58	μ = 7.17 σ = 4.36 $\tilde{x}$ = 8.35
Number of observation (N)	38	38	38	38	38

The following data was generated during the analysis of money demand equation:

ln (M/Y): The dependent variable was conducted by dividing the natural logarithm of money aggregate (ln M) to the natural logarithm of GDP (ln Y).

ln r: The natural logarithm of the interest rates was used as dependent variable in the log-log money demand specification.

DUM: a dummy variable that takes on values of 0 before the break date and values of 1 after the break date.

DUM x intercept: is merely DUM that was added to the regression where shifting in the intercept was found.

DUM x trend, DUM x r and DUM x ln r: were created by multiplying DUM with trend, r and ln r, respectively.

The following data was generated during the welfare cost estimation:

A and B: are the exponential of the intercept resulted from regression of log-log and semi-log money demand equation, respectively. In the case where the intercept is nonsignificant, the DUM x intercept is used.

2. Econometric method

The dissertation first uses a unit root test to examine the stationarity of the three variables: ln (M/Y), ln r and r, using ADF, PP and ZA tests. The null hypothesis of a non-stationary variable is tested against the alternative of stationarity. If the test statistic is lower than the critical value, then the null hypothesis is rejected, meaning that the variable is non-stationary.

Many empirical studies have used extensive versions of money demand, however within the relevance of the subject, this dissertation only specifies the demand for money in its canonical form, as such:

$$\ln(\mathrm{m}) = \ln(\mathrm{B}) - \varepsilon \mathrm{r} \tag{3.1}$$

$$\ln(\mathrm{m}) = \ln(\mathrm{A}) - \eta \ln(\mathrm{r}) \tag{3.2}$$

where B > 0 and A > 0 are the intercepts, $\varepsilon > 0$ and $\eta > 0$ are the short-term interest rates, semi-elasticity and elasticity of money demand, respectively. Real money balances are defined as money aggregate (as current LCU) divided by GDP (as constant LCU).

A structural break test is then employed to examine the stability of the two specifications (3.1) and (3.2). In particular, the one-time regime shift model by GH (1996) is utilised to test for possible shifts in the intercept or in the slope coefficients (or both). Four models are proposed by GH that are based on alternative assumptions about structural breaks: i) level shift; ii) level shift with trend; iii) regime shift where both the intercept and the slope coefficients change and iv) regime shift where intercept, trend and slope coefficients change. A single break date is endogenously determined for each of the specifications and models, such that:

GH-1: Level shift:

A simple case is that there is a level shift in the cointegrating relationship, which can be modelled as a change in the intercept μ, while the slope coefficient α is held constant. This implies that the equilibrium equation has shifted in a parallel fashion. This is called a level shift model.

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha y_{2t} + e_t \qquad t = 1, \dots, n$$

where μ_1 represents the intercept before the shift and μ_2 represents the change in the intercept at the time of the shift.

A time trend can also be added to the level shift model.

GH-2: Level shift with trend:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta t + \alpha y_{2t} + e_t \qquad t = 1, \dots, n$$

Another structural break allows for the slope to shift as well. This permits the equilibrium relation to rotate as well as to shift parallel.

GH-3: Regime shift:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha_1 y_{2t} + \alpha_2 y_{2t}\varphi_{t\tau} + e_t \qquad t = 1, \dots, n$$

In this case μ_1 and μ_2 are as in the level shift model. α_1 denotes the cointegrating slope coefficients before the regime shift, and α_2 denotes the change in the slope coefficients.

Finally, a regime shift, allowing for shifts in intercept, trend and slope coefficients, is tested for.

GH-4: Regime shift (intercept, trend and slope coefficients change):

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta_1 t + \beta_2 t\varphi_{t\tau} + \alpha_1 y_{2t} + \alpha_2 y_{2t}\varphi_{t\tau} + e_t \qquad t = 1, \dots, n$$

A conventional ADF test is applied to test for cointegration in the three alternative models above. If the null hypothesis of no cointegration is rejected, the original models cannot be used. The new equation is modified by adding dummy variables in the intercept, trend or slope variables according to the information of the break reported by the structural break tests.

Following the structural break tests, if cointegration is found, then the original money demand model cannot be used, and the modified model is used to estimate coefficient parameters using OLS regression. This step further helps identify the best money demand model for each country and for each specification by selectively choosing the equation with the most statistical significance.

Finally, the welfare cost is estimated following the consumer surplus approach proposed by Bailey (1956). The welfare cost functions for the semi-log money demand specification is expressed as:

$$W(r) = \frac{B}{\varepsilon}\left[1 - (1 + \varepsilon r)e^{-\epsilon r}\right] \qquad (3.3)$$

where B and ε are the slopes of the intercept and interest rates taken directly from equation (3.1). When the money demand takes the log-log form, then the welfare cost is:

$$W(r) = A\left(\frac{\eta}{1-\eta}\right) r^{1-\eta} \qquad (3.4)$$

where A and η are the exponential of the intercept and the elasticity of interest rates respectively, taken directly from equation (3.2). The interest rate corresponding to the steady state of the economy where inflation equals zero is assumed to be 3%. This figure has been utilised by central bankers as the target rate for long-term sustainable growth. This figure is also used in several pieces in the literature such as Lucas (2000), Ireland (2009), and Miller et al. (2014). The dissertation also estimates the welfare cost at 0%, 2% and 10% inflation that are equivalent to interest rates of 3%, 5% and 13%, respectively.

IV. Empirical results

1. Unit root test

Appendix Tables A1-E1 display results from applying ADF, PP and ZA tests for cointegration to each of the three variables: ln (M/Y), ln r and r. The tables report test statistic values of Z_t, Z(rho) and minimum t-statistic respectively for the ADF, PP and ZA unit root tests, together with the value of the lag that is either chosen manually or automatically. In particular, the ADF test is included in a trend term with 0 lag truncation; whereas the PP test is computed with a trend term and 3 lag truncation based on the Newey-West estimate and the ZA test allows for a single break in trend with the lag of 0 based on the lag method defined by Zivot-Andrews (1992). The variation in the lag truncation parameters is purposely done to increase the robustness of the test.

Insert table A1-E1 here

In all cases, the tests gave fairly unambiguous results for ln (M/Y), ln r and r. In five countries being tested, the ADF, PP and ZA tests for all three variables, in which the null hypothesis is that the variable is nonstationary, are not rejected at the 5% level. When variables are converted into first differences, the tests show no unit root with very high test statistics to indicate that the null hypothesis is rejected at the 5% level. Exceptions are found in the cases of Indonesia and Japan where variables are weakly integrated. Specifically, the unit root test report for Indonesia shows nonstationary for Δln (my) in the ZA test; whereas Japan shows weak integration for the variable Δln r with only PP Z(rho) statistic being significant at the 10% level.

In light of existing studies, the results for the unit root tests confirm the general consensus of the time-varying property of time-series data. Except for a few studies that find I(0) for interest rate variables [see Lim (1993) and Mehra (1993)], the majority of empirical work has contributed to the evidence of I(1) for series in money demand model. The same conclusion is also rendered using other testing techniques such as

KPSS (1992), Johansen (1988), Hylleberg and others (1990) or Ng and Perron (2001) [see, for example, Baba et al. (1992), Fielding (1994), Chowdhury (1995) and Miller et al. (2014)].

Since there is strong evidence that the series are integrated, it is appropriate therefore to use logarithmic models instead of linear models in depicting the long-run relationship of money demand. Furthermore, as Engle and Granger (1987) point out, before using OLS regression to estimate coefficient parameters, cointegration tests should be performed to avoid misspecification problem and misleading results. Hence, we will proceed with an analysis including the following structural break test for cointegration.

2. Structural break

Given the methodology described in the previous section, a GH test for cointegration is conducted for equations 1-4 for each of the two specifications. As shown in Appendix Tables A2-E2, the majority of the results support cointegration between variables in equations 2-4 at the 5% level. Specifically, in Indonesia, equations 3 [(break date (hereafter BD): 1995] and 4 (BD: 1998) reject the null hypothesis of no cointegration for both the semi-log and log-log specification. In Japan, only equation 3 (BD: 1999) is significant for the semi-log specification, whereas equations 1 (BD: 2003), 2 (BD: 1999) and 3 (BD: 1999) are all significant at the 5% level for the log-log specification. In Korea, equations 2 and 4 are significant in both specifications with the break dates of 1999 and 2001, respectively; cointegration is also found for equation 3 for the semi-log specification with the break date of 1988. An exception is shown in the case of the Philippines, where only equation 1 shows significance for the semi-log specification with the break date found as early as 1983. The break dates are also early in Thailand, 1982 and 1985 for equations 2 and 4, respectively, for both specifications.

Insert table A2-E2 here

The results provided by the GH test for cointegration support the existence of long-run relationships of the demand for money in the five East Asian countries. Explicitly, the results show that money demand is cointegrated with the deposit interest rate in level as well as in logarithmic form. A majority of break dates are found in the late 1990s and early 2000s, except for the Philippines and Thailand, which show break dates in the early 1980s. This is not unexpected because these countries underwent major economic reforms in the 1990s, and the break dates may highlight the importance of financial reforms in these domestic economies.

The findings are consistent with other studies. For example, Apergis (2015) applying the structural break test for cointegration from Bai and Perron (2003) found evidence of a regime shift in 1997 and 2000 for the same Asian countries. In comparison with an alternative method such as Lee and Strazicich's (2003) or Bai and Perron's (2003) structural break tests, which use endogenous break dates, the GH structural break test allows for unknown break points, which can avoid the human error that may occur.

The dissertation applies the GH method based on the premise that structural breaks may have affected the cointegrating relationships of money demand in the five East Asian countries. As incorporating econometric restrictions into estimating economic equations seems to be gaining popularity, it is necessary to test for nonlinear cointegration in addition to nonstationary tests to avoid spurious regression problems and to determine whether it is possible to model the nonstationary time-series data in an error correction model.

3. Cointegration equation

Given the presence of cointegration in the previous structural break tests, OLS regression is used to estimate the coefficient of the parameters for the equations that were cointegrated. For Indonesia, equations 3 and 4 are regressed for both money demand specifications. For Japan, equations 1-3 are regressed for the log-log specification, whereas only equation 1 is regressed for the semi-log form. For Korea, OLS regression is conducted for equations 2,3,4 and 2,4 of the semi-log and the log-log specification, respectively. For the Philippines, only equation 1 of the semi-log showed cointegration in GH test; hence, it is regressed. For Thailand, equations 2 and 4 of both specifications are regressed by the OLS.

This step further helps identify the best equation as well as the break date by considering the significance of these coefficients. As reported in Appendix Tables A3-E3, results indicate the best break date: for Indonesia – 1998, Japan – 2003, Korea – 1999, Philippines – 1983 and Thailand – 1982.

Insert table A3-E3 here

These break dates are sensible as they are often associated with major economic events that occurred in each country. Specifically, the break date identified for Indonesia in 1998 is associated with the financial crisis occurring in the country and other Asian economies during this period. The financial crisis had caused the domestic currency to devalue almost 75% from 1997 to 1998. Inflation was climbing at a rate of 20% per year. During this period, the government experienced a series of financial reforms influenced by the International Monetary Fund (IMF) rescue package. (Sherlock, 1998)

In Japan, after a decade of "great recession" from 1990 to early 2000s as falling GDP reached its peak in 2003, the economy started to recover with support from specialised exports that helped to boot up investment, while the Nikkei index recovered and broke its 10-months record. The inflation rate, which had been near zero, started to increase into the positive, demonstrating the economy escaping the deflationary state.

Similarly, to Japan, the Korean economy had been affected by the global financial crisis in 1997; however, the country quickly recovered its economic position with a series of reform programs implemented under the newly elected president Kim Dae-jung in 1999.

The Philippines, on the other hand, experienced major changes in economic structure and fiscal policies in 1983, earlier than the other countries. The increase in government expenditure through money supply consequently depreciated the peso three times in one year between June 1983 and June 1984. The negative growth of output (-5.9% in 1983) along with the climbing rates of inflation has given the Philippines the most elastic interest rate (0.05) among the five countries. (Bucog, 2004)

Thailand's economy also experienced early structural changes in 1982 that were largely impacted by the deceleration of prices worldwide, which reduced domestic inflation rates from 13% in 1981 to only 5% in 1982. Further, public expenditures also experienced a significant drop from 64.7% in 1980 to 8.4% in 1982, due to reduced public saving. (LePoer, 1987)

The interest rate elasticities estimated for the five countries fall in the range of 0.002 to 0.018, whereas the interest rate semi-elasticities fall in the range of 0.0004 to 0.05 and are all negative in sign. These estimates are relatively insignificant, indicating an irresponsive effect of interest rates to money demand. Our results are smaller than those estimated for developed countries such as the US [of 0.123 – Miller et al. (2014)], EU members [of 0.1-0.52 – Serletis and Yavari (2007)] and also for developing countries [of 0.08-0.306 – see James (2005), Atkins (2005), and Nair et al. (2008)]. In a comparison between the two specifications, except for the Philippines and Japan, the rest of the countries report an interest rate elasticity of approximately ten times higher than the interest rate semi-elasticity. Overall, the magnitude and the sign of the interest rate elasticities adhere to economic theory and are closer to the empirical evidence for developing countries.

The intercepts estimated for the five countries lie between 2.43 and 5.34 (in absolute value) except for the Philippines and Japan. Intuitively, these figures indicate the average value of $\ln\left(\frac{M}{Y}\right) + \eta\ln(r)$ if the equation takes on the log-log form or the average value of $ln\ \left(\frac{M}{Y}\right) + \ \varepsilon r$ if the equation takes on the semi-log form. In comparison to other studies such as Akinlo (2006), who reported a value of 3.52 for the intercept, our estimates are justifiable.

Overall, the slope coefficients estimated by OLS regression for the five East Asian countries achieve the expected signs and are statistically significant at the 95% confidence level. Moreover, the log-log form is more consistent than the semi-log, though not significant, in estimating the money demand model.

4. Welfare estimate

After collecting results from OLS regression, the best equation, which is the most statistically significant among the options, is picked for each specification to estimate the welfare cost. Following the assumption made by Lucas (2000) that the steady-state real interest rate equals 3%, so that r = 0.03 corresponds to zero inflation, r = 0.05 corresponds to 2% annual inflation, and r = 0.13 corresponds to 10% annual inflation, the welfare cost is computed based on equation (2.13) and (2.14) for the log-log and semi-log specification, respectively. Table 2 and Table 3 present the welfare cost estimates for the five East Asian countries for the semi-log and log-log specifications, respectively.

Table 2: Welfare cost estimation (semi-log)

Country	Estimated equation		Estimated welfare (percent of income)		
	B=exp(α)	**ε=β**	**0% inflation W(0.03)**	**2% inflation W(0.05)**	**10% inflation W(0.13)**
Indonesia	0.0052	0.0004⁺	$9.36*10^{-8}$	$2.60*10^{-7}$	$1.76*10^{-6}$
Japan					
Korea	0.075	0.001	$3.37*10^{-6}$	$9.37*10^{-6}$	$6.34*10^{-5}$
Philippines	2.4351	0.05	0.00547	0.01519	0.10244
Thailand	0.0883	0.0004	$1.59*10^{-6}$	$4.41*10^{-6}$	$2.98*10^{-5}$

⁺ denotes the slope coefficient of the regime dummy variable

Table 3: Welfare cost estimation (log-log)

Country	Estimated equation		Estimated welfare (percent of income)		
	A=exp(α)	η=β	0% inflation W(0.03)	2% inflation W(0.05)	10% inflation W(0.13)
Indonesia	0.0048	0.007⁺	$1.04*10^{-4}$	$1.73*10^{-4}$	$4.46*10^{-4}$
Japan	2.6459	0.006	0.05	0.08	0.21
Korea	0.078	0.018	0.005	0.008	0.019
Philippines					
Thailand	0.0853	0.002	$5.16*10^{-4}$	$8.60*10^{-4}$	$2.23*10^{-3}$

⁺ denotes the slope coefficient of the regime dummy variable

To provide a comparison of welfare cost estimates across the different countries, we assumed a 10% increase in interest rates (from 3% to 13%). Table shows that increasing the interest rate by this amount for the semi-log money demand would yield a cost that is equivalent to a reduction in real income of $1.77*10^{-6}$ % in Indonesia,

$4.97*10^{-5}$ % in Korea, 0.10% in the Philippines, and $2.98*10^{-5}$ % in Thailand. For the log-log money demand, Table 5 shows a cost of $4.50*10^{-4}$ % in Indonesia, 0.21% in Japan, 0.061% in Korea, and 0.002% in Thailand.

Several observations can be derived from the results in Table 2 and Table 3. First, the distribution of the welfare cost of inflation is heterogeneous across the five East Asian countries. Second, on average in the countries studied, 10% inflation is worth less than a half percent of consumption. The average welfare cost of 10% inflation for the five countries is 0.025 and 0.068% of income for the semi-log and the log-log specifications, respectively. In only two countries does the welfare cost exceed 0.1%. The Philippines and Japan, two countries for which the model exhibits a good fit to the data, in general, have welfare costs of 0.1% and 0.21%, respectively. Indonesia is at the opposite end of the spectrum, with less than 0.01%. Overall, these findings are lower than the welfare cost estimates for the US of 0.2-0.45% [see Dotsey and Ireland (1994), Ireland (2009) and Miller et al. (2014)], and for European countries of 0.1-0.4% [see Serletis and Yavari (2007)], and much lower than the estimates for African countries of 0.9-1.08% [see Makochekanwa (2008) and Gupta and Uwilingiye (2008)].

When the interest rate reduces from 3% to 5%, it would yield a cost equivalent to a reduction in real income of on average 0.002% and 0.007% for the semi-log and log-log specifications, respectively. These figures are in line with Serletis and Yavari (2004) but much smaller than Lucas (2000) calculated for the US during 1900–1994. Moreover, increasing the interest rate from 5% to 13% would yield a cost equivalent to 0.016% and 0.034% of real income on average. These are, however, larger than the 0.0045% obtained by Serletis and Yavari (2004), but still smaller than the 0.9% from Lucas (2000). For the Philippines and Japan, the changes are significantly higher than for other countries; increasing the interest rate from 3% to 5% would reduce real income by 0.01% and 0.03%, whereas increasing the interest rate from 5% to 13% would reduce real income by 0.08% and 0.13%, respectively.

Why are our estimates of the welfare cost of inflation significantly lower than those reported by other studies? To answer this question, we re-examined the results that have been estimated by this dissertation. We revealed that the major cause of the issue comes from the inelasticity of interest rates estimated. Following Bailey's consumer surplus approach, interest rate elasticity is an important input in the welfare estimation function. In comparison with studies where the interest rate elasticity is endogenously set around unity, for example Lucas (2000) prefers this rate at 0.5, our estimates are far below this figure. There is still a big gap among studies using a similar econometric approach. For example, Makochekanwa (2008) estimates the interest elasticity as 0.247 and 2.019 respectively for the log-log and semi-log specifications. These figures are much larger than our estimates of 0.004 and 0.013 for the log-log and semi-log. We conclude therefore that estimation of the interest elasticity of money demand is crucial in evaluating the welfare cost of inflation.

Considering another aspect, there is a difference between the semi-log and the log-log specification in the order of magnitude of the welfare estimated. Except for the Philippines and Japan, results obtained from the semi-log specification are significantly low, falling in the range of $4.97*10^{-5}$ to $1.77*10^{-6}$ % of income at 10% inflation, while the log-log specification seems to provide more plausible estimates that lie between 0.061 and $4.50*10^{-4}$ % of income. This finding is in line with the value Makochekanwa (2008) estimated for Zimbabwe during 1990–2005 using monthly data, however, it is opposite to his estimates for the period 1980–2005 using quarterly data. This is again due to the issue of interest rate elasticity estimate. To this extent, there is a lack of empirical evidence to justify the difference between the two specifications. Ireland (2009) is one of a few who estimated the money demand relationship for both specifications, in which he obtains similar results with the semi-log yielding a higher interest rate elasticity (of 1.794) compared to the log-log (of 0.087).

V. Conclusion

This dissertation revisits the estimation of the welfare cost of inflation incorporating time-varying tests and structural break tests for cointegration with the log-log and semi-log specifications for five selected Asian countries, i.e., Indonesia, Japan, Korea, the Philippines and Thailand, from 1977 to 2014.

In unit root tests, results indicate strong evidence of integration of first order for all of the variables in both specifications. This evidence enables us to proceed with other tests and estimations as variables are cointegrated according to the Engle-Granger (1987) representation theorem. Moreover, the use of alternative testing techniques (such as PP and ZA) and procedures (such as allowing for a break in the trend) have increased the robustness of the evidence of nonstationary variables.

The structural break tests also show cointegration with different break dates for different types of shifts in each country. In particular, Indonesia shows a break date in 1995 for the GH3 type and in 1998 for the GH4 type; Japan shows a break date in 2003 for the GH1, GH2, and GH3 types and in 1999 for the GH4 type; Korea shows a break date in 1999 for the GH2 type, in 1988 for the GH3 type, and in 2001 for the GH4 type; the Philippines shows a break date in 1983 for the GH1 type; and Thailand shows a break date in 1982 for the GH2 type and in 1985 for the GH4 type.

The main test to estimate the money demand function is OLS, which further helps to identify the best model and break date for each country based on the statistical significance of the coefficients. Moreover, the break dates identified by the tests are sensibly taking into account historical economic events that occurred in each country.

Although this dissertation is not able to pinpoint the proper specification of the money demand function with respect to the form of the interest rate, the log-log specification seems to provide a better fit in comparison with the semi-log form. In general, the estimates of interest elasticity conform to the theoretical expectation and the empirical evidence from the previous literature, lying between 0.05 and 0.0004 (in absolute value) with a negative sign. Moreover, as Ireland (2009) pointed out, the key issue is not so much whether the demand for money depends on the logarithm or the level of the nominal interest rate, but instead whether there exists some finite satiation point that places a limit on money demand when expressed as a fraction of real income.

The welfare cost of inflation estimated by the dissertation is an order of magnitude smaller than those obtained from other empirical studies. While Japan and the Philippines deviate from the mean with a significantly higher cost of inflation (0.21% and 0.1% of income at the 10% inflation), other countries experience a much lower cost, between 0.06 and 0.0004% of income for the log-log specification and between $2.98*10^{-5}$ and $1.77*10^{-6}$ % of income for the semi-log specification. These estimates are lower than those estimated for the US of 0.2-0.4% and lower than for European countries of 0.1-0.4%. This observation suggests a new consideration for the level of the cost of inflation in Asian region. However, it is worthwhile to note that these welfare cost estimates account for only the money demand distortion brought about by positive nominal interest rates. Dotsey and Ireland (1996) demonstrate that increases in inflation can have an impact on other marginal decisions such as aggregate output in level as well as in growth rate, while Martin Feldstein (1997) argues that the interactions between inflation and a tax code that is not completely indexed can add substantially to the welfare cost of inflation.

The dissertation is not without limitations. First, the sparsity of the data with only 38 observations of annual data for each country, due to availability issues, has prevented the model from rendering accurate estimations of the money demand relationship. Second, the long-run relationship between the money aggregate and the interest rate can be investigate further using Granger causality test to check for the direction of the impact between variables.

Finally, the overall results suggest a low cost of inflation under the contemporary financial system which has been moving from a money aggregate target to an interest rate base. However, as it is argued by Serletis and Yavari (2007), much of the welfare cost of inflation is borne by the poor and thus impact directly to the income distribution, and cannot accurately be assessed using aggregate methods.

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VII. Appendix

Tabulate for Indonesia

Table A1: ADF, PP and ZA unit root test, 1977–2014

Variables	LAG	ADF	LAG	Philips-Perron	LAG	Zivot-Andrews
ln (M/Y)	[0,0]	-1.275	[3,0]	-7.575	[0,0]	-2.597
		(3.552)		(18.812)		(4.42)
Δ ln (M/Y)	[0,1]	-3.685	[3,1]	-20.514	[0,1]	-4.085
		(3.556)*		(18.736)*		(4.42)
r	[0,0]	-2.405	[3,0]	-9.596	[1,0]	-3.011
		(3.552)		(18.812)		(4.42)
Δ r	[0,1]	-6.115	[3,1]	-30.644	[0,1]	-6.479
		(3.556)*		(18.736)*		(4.42)*
ln r	[0,0]	-1.916	[3,0]	-6.534	[0,0]	-3.472
		(3.552)		(18.812)		(4.42)
Δ ln r	[0,1]	-5.810	[3,1]	-29.284	[0,1]	-6.314
		(3.556)*		(18.736)*		(4.42)*

The LAG columns indicate number of lags and order of the variable.
* and ** denote significance at the 5 and 10% level, respectively.

Table A2: Structural break test for cointegration, 1977–2014

Specification/ GH model	Break date	GH test statistic	5% critical value	Existence of cointegration
$\ln(m) = \ln(B) - \varepsilon r$				
GH-1	2000	-3.95	(4.61)	No
GH-2	1999	-4.11	(4.99)	No
GH-3	1995	-5.65	(4.95)	Yes
GH-4	1998	-5.50	(5.50)	Yes
$\ln(m) = \ln(A) - \eta \ln(r)$				
GH-1	2000	-3.65	(4.61)	No
GH-2	1998	-4.14	(4.99)	No
GH-3	1995	-5.60	(4.95)	Yes
GH-4	1998	-5.16	(5.50)	Yes

Table A3: Cointegration equation, 1977–2014

	Specification 1				Specification 2			
	GH-1	**GH-2**	**GH-3**	**GH-4**	**GH-1**	**GH-2**	**GH-3**	**GH-4**
Intercept			-5.843 (13.07) *	-5.251 (15.04)*			-5.822 (11.57)*	-.5.331 (14.26)*
DUM x Intercept			.005 (0.62)	1.552 (2.69)*			.007 (0.39)	1.755 (2.59)*
Trend			.003 (14.91)*	.003 (17.39)*			.003 (13.14)*	.003 (16.36)*
DUM x Trend				-.001 (2.65)*				-.001 (2.55)*
r			-.0002 (0.83)	-.0001 (0.28)				
DUM x r			-.00001 (0.04)	-.0004 (1.94) **				
ln r							-003 (0.86)	-.001 (0.54)
DUM x ln r							-.0005 (0.09)	-.007 (1.93)**

* and ** denote significance at the 5 and 10% level, respectively.

Tabulate for Japan

Table B1: ADF, PP and ZA unit root test, 1977–2014

Variables	LAG	ADF	LAG	Philips-Perron	LAG	Zivot-Andrews
ln (M/Y)	[0,0]	-1.686	[3,0]	-4.766	[1,0]	-2.282
		(3.552)		(18.812)		(4.42)
Δ ln (M/Y)	[0,1]	-4.540	[3,1]	-29.466	[0,1]	-5.684
		(3.556)*		(18.736)*		(4.42)*
r	[0,0]	-2.183	[3,0]	-11.802	[1,0]	-4.975
		(3.552)		(18.812)		(4.42)
Δ r	[0,1]	-4.513	[3,1]	-25.947	[0,1]	-4.956
		(3.556)*		(18.736)*		(4.42)*
ln r	[0,0]	-1.093	[3,0]	-6.183	[1,0]	-3.618
		(3.552)		(18.812)		(4.42)
Δ ln r	[0,1]	-3.177	[3,1]	-17.534	[0,1]	-3.398
		(3.556)		(18.736)**		(4.42)

The LAG columns indicate number of lags and order of the variable.
* and ** denote the rejection of the null at the 5 and 10% level, respectively.

Table B2: Structural break test for cointegration, 1977–2014

Specification/ GH model	Break date	GH test statistic	5% critical value	Existence of cointegration
$\ln(m) = \ln(B) - \varepsilon r$				
GH-1	2004	-3.77	(4.61)	No
GH-2	2000	-3.15	(4.99)	No
GH-3	1999	-4.85	(4.95)**	Yes
GH-4	1998	-3.96	(5.50)	No
$\ln(m) = \ln(A) - \eta \ln(r)$				
GH-1	2003	-5.25	(4.61)	Yes
GH-2	2003	-5.00	(4.99)	Yes
GH-3	1999	-5.15	(4.95)	Yes
GH-4	1999	-4.52	(5.50)	No

* and ** denote significance at the 5 and 10% level, respectively.

Table B3: Cointegration equation, 1977–2014

	Specification 1				Specification 2			
	GH-1	GH-2	GH-3	GH-4	GH-1	GH-2	GH-3	GH-4
Intercept			-.285 (1.22)		.973 (1620.66)*	.014 (0.08)	.126 (0.62)	
DUM x Intercept			.011 (5.61)*		.023 (19.73)*	.016 (9.70)*	.021 (8.94)*	
Trend			.001 (5.41)*			.000 (5.52)*	.000 (4.17)*	
DUM x Trend							.006 (5.27)*	
r			-.002 (2.74)*					
DUM x r			.006 (1.79)**					
ln r					-.006 (15.79)*	-.004 (9.20)*	-.004 (5.62)*	
DUM x ln r								

* and ** denote the rejection of the null at the 5 and 10% level, respectively.

Tabulate for Korea

Table C1: ADF, PP and ZA unit root test, 1977–2014

Variables	LAG	ADF	LAG	Philips-Perron	LAG	Zivot-Andrews
ln (M/Y)	[0,0]	-2.427	[3,0]	-12.599	[0,0]	-2.585
		(3.552)		(18.812)		(4.42)
Δ ln (M/Y)	[0,1]	-6.250	[3,1]	-38.362	[0,1]	-6.416
		(3.556)*		(18.736)*		(4.42)*
r	[0,0]	-2.848	[3,0]	-16.262	[1,0]	-4.368
		(3.552)		(18.812)		(4.42)
Δ r	[0,1]	-5.759	[3,1]	-29.028	[0,1]	-5.764
		(3.556)*		(18.736)*		(4.42)*
ln r	[0,0]	-2.762	[3,0]	-15.898	[0,0]	-3.070
		(3.552)		(18.812)		(4.42)
Δ ln r	[0,1]	-5.834	[3,1]	-31.479	[0,1]	-5.864
		(3.556)*		(18.736)*		(4.42)*

The LAG columns indicate number of lags and order of the variable.
* and ** denote significance at the 5 and 10% level, respectively.

Table C2: Structural break test for cointegration, 1977–2014

Specification/ GH model	Break date	GH test statistic	5% critical value	Existence of cointegration
$\ln(m) = \ln(B) - \varepsilon r$				
GH-1	1999	-4.27	(4.61)	No
GH-2	1999	-5.41	(4.99)	Yes
GH-3	1988	-5.78	(4.95)	Yes
GH-4	2001	- 6.19	(5.50)	Yes
$\ln(m) = \ln(A) - \eta \ln(r)$				
GH-1	1998	-4.14	(4.61)	No
GH-2	1999	-5.19	(4.99)	Yes
GH-3	1989	-4.57	(4.95)	No
GH-4	2001	-6.20	(5.50)	Yes

Table C3: Cointegration equation, 1977–2014

	Specification 1				Specification 2			
	GH-1	GH-2	GH-3	GH-4	GH-1	GH-2	GH-3	GH-4
Intercept		-2.833 (5.06)*	-2.74 (4.46)	-2.59 (12.34)*		-2.393 (4.47)*		-2.555 (12.47)*
DUM x Intercept		.018 (3.57)*	.043 (3.87)*	2.071 (3.88)*		.014 (2.91)*		2.097 (3.90)*
Trend		.002 (6.73)*	.002 (5.94)*	.002 (16.78)*		.002 (6.32)*		.002 (17.28)*
DUM x Trend				-.001 (3.84)*				-.001 (3.90)*
r		-.001 (2.24)*	-.001 (1.09)*	-.001 (3.47)*				
DUM x r			-.004 (4.97)*	-.001 (0.73)				
ln r						-.018 (3.48)*		-.009 (3.80)*
DUM x ln r								.003 (0.58)

* and ** denote significance at the 5 and 10% level, respectively.

Tabulate for Philippines

Table D1: ADF, PP and ZA unit root test, 1977–2014

Variables	LAG	ADF	LAG	Philips-Perron	LAG	Zivot-Andrews
ln (M/Y)	[0,0]	-1.273	[3,0]	-3.2	[0,0]	-4.733
		(3.552)		(18.812)		(4.42)
Δ ln (M/Y)	[0,1]	-7.766	[3,1]	-43.661	[0,1]	-8.546
		(3.556)*		(18.736)*		(4.42)*
r	[0,0]	-3.072	[3,0]	-11.615	[1,0]	-6.050
		(3.552)		(18.812)		(4.42)
Δ r	[0,1]	-5.105	[3,1]	-24.891	[0,1]	-5.705
		(3.556)*		(18.736)*		(4.42)*
ln r	[0,0]	-1.963	[3,0]	-5.965	[0,0]	-3.365
		(3.552)		(18.812)		(4.42)
Δ ln r	[0,1]	-6.207	[3,1]	-31.779	[0,1]	-5.197
		(3.556)*		(18.736)*		(4.42)*

The LAG columns indicate number of lags and order of the variable.
* and ** denote the rejection of the null at the 5 and 10% level, respectively.

Table D2: Structural break test for cointegration, 1977–2014

Specification/ GH model	Break date	GH test statistic	5% critical value	Existence of cointegration
$\ln(m) = \ln(B) - \varepsilon r$				
GH-1	1983	-4.94	(4.61)	Yes
GH-2	1987	-3.68	(4.99)	No
GH-3	1987	-4.67	(4.95)	No
GH-4	2000	-5.09	(5.50)	No
$\ln(m) = \ln(A) - \eta \ln(r)$				
GH-1	1985	-3.93	(4.61)	No
GH-2	1987	-4.04	(4.99)	No
GH-3	1987	-4.16	(4.95)	No
GH-4	2000	-4.79	(5.50)	No

Table D3: Cointegration equation, 1977–2014

	Specification 1				Specification 2			
	GH-1	GH-2	GH-3	GH-4	GH-1	GH-2	GH-3	GH-4
Intercept	.89 (93.01)*							
DUM x Intercept	.068 (8.55)*							
Trend								
DUM x Trend								
r	-.05 (8.99)*							
DUM x r								
ln r								
DUM x ln r								

* and ** denote significance at the 5 and 10% level, respectively.

Tabulate for Thailand

Table E1: ADF, PP and ZA unit root test, 1977–2014

The LAG columns indicate number of lags and order of the variable.

Variables	LAG	ADF	LAG	Philips-Perron	LAG	Zivot-Andrews
ln (M/Y)	[0,0]	-3.341	[3,0]	-18.220	[0,0]	-4.758
		(3.552)		(18.812)		(4.42)
Δ ln (M/Y)	[0,1]	-7.529	[3,1]	-39.489	[0,1]	-5.374
		(3.556)*		(18.736)*		(4.42)*
r	[0,0]	-2.785	[3,0]	-11.118	[1,0]	-2.823
		(3.552)		(18.812)		(4.42)
Δ r	[0,1]	-5.500	[3,1]	-31.904	[0,1]	-5.839
		(3.556)*		(18.736)*		(4.42)*
ln r	[0,0]	-2.486	[3,0]	-10.735	[0,0]	-3.665
		(3.552)		(18.812)		(4.42)
Δ ln r	[0,1]	-4.606	[3,1]	-24.135	[0,1]	-4.735
		(3.556)*		(18.736)*		(4.42)*

* and ** denote the rejection of the null at the 5 and 10% level, respectively.

Table E2: Structural break test for cointegration, 1977–2014

Specification/ GH model	Break date	GH test statistic	5% critical value	Existence of cointegration
$\ln(m) = \ln(B) - \varepsilon r$				
GH-1	2008	-3.46	(4.61)	No
GH-2	1982	-5.56	(4.99)	Yes
GH-3	2007	-3.51	(4.95)	No
GH-4	1985	-6.01	(5.50)	Yes
$\ln(m) = \ln(A) - \eta \ln(r)$				
GH-1	2007	-3.68	(4.61)	No
GH-2	1982	-5.44	(4.99)	Yes
GH-3	1990	-3.46	(4.95)	No
GH-4	1985	-5.89	(5.50)	Yes

Table E3: Cointegration equation, 1977–2014

	Specification 1				Specification 2			
	GH-1	GH-2	GH-3	GH-4	GH-1	GH-2	GH-3	GH-4
Intercept		-2.427 (16.38)*		-5.161 (1.82)*		-2.461 (16.63)*		-.124 (0.08)
DUM x Intercept		-.003 (2.31)*		2.867 (1.01)		-.003 (2.34)*		-2.377 (1.61)**
Trend		.002 (22.71)*		-.003 (2.14)*		.002 (23.00)*		.0005 (0.68)
DUM x Trend				-.001 (1.01)				.001 (1.61)**
r		-0.0004 (2.78)*		-0005 (0.47)				
DUM x r				-.0001 (0.12)				
ln r						-.002 (3.06)*		.008 (1.07)
DUM x ln r								-.010 (1.33)**

* and ** denote significance at the 5 and 10% level, respectively.