MONEY DEMAND AND THE WELFARE COST OF INFLATION: EMPIRICAL EVIDENCE FROM EAST ASIAN COUNTRIES

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Abstract: This article 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.
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1. 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). The main contribution of this research is to provide evidence of the welfare cost of inflation in Asian region using Bailey's (1956) method. The five selected countries include Indonesia, Japan, Korea, Philippines and Thailand. These countries are among the most stable economies in Asia where the markets are freely behaved.

The study sheds additional light on this issue by incorporating a structural break technique in the computation process. To the author's best knowledge, there have not yet been any studies estimating the welfare cost based on structural break analysis. The use of this research method is to examine the stability of the model and further to reduces estimation error occurred by possible shift in the intercept vector or the slope coefficient vectors (or both).

2. Literature review

2.1. Theoretical framework

The modern theory of money demand is illuminated by the Keynes (1936) theory, which 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

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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: $m_t = c_t - \alpha r_t - \alpha \pi_t^e$ where mt = 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: $\ln\left(\frac{M}{P}\right) = \alpha + \beta lnY + \gamma R$

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

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. [13] A reflection of the short-run demand for money function can be obtained as: $\log\left(\frac{M_t}{P_t}\right) = \lambda\alpha + \lambda\beta_1 logY_t + \lambda\beta_2 R_t + (1-\lambda)\ log\left(\frac{M_{t-1}}{P_{t-1}}\right) + e_t$ 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. 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. 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. 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: $\Delta y_t = \alpha_1 y_{t-1} + e_t$ where Δyt 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: $\Delta y_t = \sum_{j=1}^{p} a_j y_{t-j} + e_t$

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 DUt, DTt and $D(T_B)_t$ to the model: $y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_B)_t + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + e_t$ 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. This article will utilize the structural break technique by GH (1996) to examine the stability and to estimate the money demand specification for each country. Further illustration of this method will be presented in the following section of the article.

2.2. Empirical Evidence

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.

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.

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.

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.

There is a lack of literature estimating the welfare cost of inflation in Asian region, especially in East-Asia region. Recognising the fact, this article is crucial in fulfilling our knowledge and understanding of the relationship between inflation and welfare.

3. Data and methodology

3.1. Data

This research 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, 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 research 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.

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.

3.2. Econometric model

The study 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 research only specifies the demand for money in its canonical form, as such:

$$\ln(\mathrm{m}) = \ln(\mathrm{B}) - \varepsilon \mathrm{r} \qquad (5)$$

$$\ln(\mathrm{m}) = \ln(\mathrm{A}) - \eta \ln(\mathrm{r}) \qquad (6)$$

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 (5) and (6). In particular, the one-time regime shift model by GH (1996) is utilised to test for possible shifts either 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:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha y_{2t} + e_t \qquad (7.1)$$
$$t = 1, \dots, n$$

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

GH-2: Level shift with trend:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha y_{2t} + e_t \qquad (7.2)$$
$$t = 1, \dots, n$$

GH-3: Regime shift:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1 y_{2t} + \alpha_2 y_{2t} \varphi_{t\tau} + e_t \qquad (7.3)$$
$$t = 1, \dots, n$$

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.

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

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta_1 t + \beta_2 t \varphi_{t\tau} + \alpha_1 y_{2t} + \alpha_2 y_{2t} \varphi_{t\tau} + e_t \qquad t = 1, \dots, n \ (7.4)$$

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. When the money demand takes on log-log and semi-log form, equation (3) and (4) are applied as the welfare cost function, respectively.

The interest rate corresponding to the steady state of the economy where inflation equals zero is assumed to be 3%. This figure has been widely utilised by central bankers as the target rate for long-term sustainable growth as well as in various literatures. [7], [9], [11] The research also estimates the welfare cost at 0%, 2% and 10% inflation that are equivalent to interest rates of 3%, 5% and 13%, respectively.

4. Research results

4.1.Unit root tests

In all cases, unit root 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. (Luan, 2016)

4.2. Structural breaks test

Given the methodology described previously, a GH test for cointegration is conducted for equations (7.1)-(7.4) for each of the two specifications. 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. [8]

4.3. Welfare estimate

After collecting results from OLS regression, the best equation, which is the most statistically significant among the options, was selected for each specification to estimate the welfare cost. The welfare cost is computed based on equation (3) and (4) for the log-log and semi-log specification, 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	$5.16*10^{-4}$	$8.60*10^{-4}$	$2.23*10^{-3}$

Results show that increasing the interest rate by 10% (from 3% to 13%) 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, these costs are equivalent to a reduction in real income of $4.50*10^{-4}$ % in Indonesia, 0.21% in Japan, 0.061% in Korea, and 0.002% in Thailand.

5. Conclusion

This study 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.

Although the study is not able to pinpoint a proper specification for 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. (Ireland, 2009)

The welfare cost of inflation estimated by the research is an order of magnitude smaller than those obtained from other empirical studies. 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 welfares 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.

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. (Serletis, A., & Yavari, K., 2007)

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