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Abstract

The emerging consensus in monetary policy and business cycle analysis is that money aggregates are not useful as an intermediate target for monetary policy or as an information variable. The uselessness of money as an intermediate target is driven by empirical research that suggests that money demand is unstable. In addition, the informational quality of money has been called into question by empirical research that fails to identify a relationship between money growth and inflation, nominal income growth, and the output gap. Nevertheless, this research is potentially flawed by the use of simple sum money aggregates, which are not consistent with economic, aggregation, or index number theory. This paper therefore re-examines previous empirical evidence on money demand and the role of money as an information variable using monetary services indexes as monetary aggregates. These aggregates have the advantage of being derived from microtheoretic foundations as well as being consistent with aggregation and index number theory. The results of the re-evaluation suggest that previous empirical work might be driven by mismeasurement.

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1 Introduction

The emerging consensus in the literature on monetary policy and business cycles is that money aggregates can be altogether ignored without the loss of significant information. The justification for this consensus is based on four factors. First, the Federal Reserve and other central banks around the world use an interest rate as their monetary policy instrument. As such it is not clear a priori whether money aggregates provide additional information not communicated by movements in the interest rate. Second, there is a widespread belief that the demand for money is unstable (Friedman and Kuttner, 1992; Estrella and Mishkin, 1997; Woodford, 1998) and as a result money aggregates do not have a predictable influence on other economic variables. Third, empirical estimation of backward-looking IS equations do not find a statistically significant relationship between real money balances and the output gap (Rudebusch and Svensson, 2002). Finally, the dynamic New Keynesian model, which is the workhorse of modern monetary policy research, abstracts from money completely based on the claim that money is redundant in the model.[2] In fact, McCallum (2001) notes that the quantitative implications of this omission are quite small.

While the empirical research cited above casts doubts on the role of money serving as either an information variable or an intermediate target for monetary policy, this evidence is potentially flawed by the use of simple sum money aggregates. The use of simple sum aggregates is problematic because this aggregation procedure is only valid in the case in which all money assets in the particular aggregate are perfect substitutes. This limiting case is not supported empirically.

An alternative to the simple sum aggregates is the monetary services index, first derived by Barnett (1980) and available through the St. Louis Federal Reserve FRED database.[3] The advantage of using the monetary services index is that the index is derived from mi-

[2] For a textbook treatment, see Woodford (2003) or Gali (2008).

[3] There exist corresponding indexes for M1, M2, M3, and MZM. The database also includes the currency equivalent aggregates developed by Rotemberg, Driscoll, and Poterba (1995). These latter aggregates are not used in this paper.

crotheoretic foundations and is consistent with aggregation and index number theory.

The purpose of this paper is to re-examine results that have been used to justify the exclusion of money using the monetary services indexes rather than the simple sum money aggregates. The paper therefore posits the claim that these earlier results are flawed by mismeasurement. The paper proceeds as follows. Section 2 introduces the dynamic New Keynesian model and the corresponding theoretical arguments for the exclusion of money as well as a brief discussion of the supporting empirical results. Second 3 examines the existence and significance of mismeasurement for monetary aggregates. Section 4 provides empirical evidence on money demand and the IS equation with and without money. Section 5 concludes.

2 The Cashless Approach

2.1 Theory

The baseline New Keynesian model captures the cashless approach rather concisely. Combined with an interest rate rule, equilibrium can be defined by the following equations.

$$\tilde{y}_t = E_t\tilde{y}_{t+1} - (R_t - E_t\pi_{t+1}) + \varepsilon_t^{IS} \tag{1}$$

$$\pi_t = \beta E_t\pi_{t+1} + \kappa\tilde{y} \tag{2}$$

where $\tilde{y}$ is the output gap, π is the inflation rate, and R_t is the nominal interest rate. Equation (1) is the dynamic IS equation and equation (2) is the New Keynesian Phillips curve.

The exclusion of money from the model is the result of the purported instability of money demand and the lack of a role for money as an information variable independent of the interest rate. As Woodford (1998) explains, models that use the path of money to predict inflation and other economic variables assume a stable money demand function. More specifically, the equilibrium price level is determined as the level of prices that equate the purchasing power of the existing money supply with the demand for real balances. However, if the demand for money is not stable as the evidence in the subsection below suggests, it is potentially useful to identify an alternative formulation and explanation of monetary policy. The cashless approach is one such method as it enables one to determine the price level without reference to the money supply. The dynamic New Keynesian model above does so by replacing a traditional LM-type relation with a rule that describes the path of the interest rate.

The cashless approach, however, takes the analysis beyond the role of money demand stability. By explicitly excluding money from the model, the cashless approach necessarily implies that there is no informational content in the behavior of the money supply or real money balances beyond what is already reflected in the interest rate. In other words, money is redundant.

Typically, following Woodford (2003), the exclusion of money is justified by the fact that the inclusion of real money balances as a separable argument of the utility function is fundamentally equivalent to the cashless approach. In addition, one can derive an IS equation with real money balances by modeling real balances as non-separable with regards to consumption in the utility function. However, under reasonable parameterizations of the model, such a real balance effect would be small.

While these latter arguments are theoretically valid, they ignore a key insight of monetarist thought with regards to the behavior of real money balances. The real balance effect is typically thought of as a wealth effect that arises from a change in nominal money balances. Traditionally, however, this is not the effect that monetarists emphasized with regard to real money balances. Rather, monetarists traditionally view real money balances as a sort of index that captures the variety of relative price effects across a broad spectrum of assets. This view underlies the belief that monetary transmission mechanism varies in terms of the direction of the flow of funds and the lag in the effect of policy. As such, the behavior of real money balances are able to reflect the monetary transmission process.[4] It is therefore possible that the cashless approach excludes assets that are important to the monetary transmission mechanism and that these effects could be communicated through the behavior

[4]For a summary of this view, see Meltzer (2001) and Nelson (2003).

of real money balances if the money demand function was richly specified in the model.

Ultimately, the cashless approach is built upon the existence of instability in money demand and the lack of a unique, informational role for money. These represent clear empirical hypotheses that can be tested. The stability of money demand and the role of money as an information variable are explored below.

2.2 Evidence

It is by now a well-accepted axiom that a stable money demand function is a necessary condition for money to exert a predictable influence on economic variables. What's more, there exists an emerging consensus in the literature that money demand has been unstable since the beginning of the 1980s and that money is not useful as an information variable. Indeed, this is a primary justification for the cashless approach outlined above. Specifically, the work of Friedman and Kuttner (1992) and Estrella and Mishkin (1997) are often cited as providing comprehensive evidence of this view. These results are discussed in turn below.

Friedman and Kuttner (1992) conduct a comprehensive analysis of money demand stability and the role of money as an information variable by employing two broad approaches. First, they examine the role of money growth in influencing nominal income growth and inflation under the assumption that if money is useful as an information variable, it should be the primary predictor of each. The second approach is to measure the stability of money demand using a cointegrated vector autoregressive (VAR) approach. For each of their approaches, they estimate results for three samples, the first sample runs from 1960:2 - 1979:3, the second from 1960:2 - 1990:4, and the final from 1970:3 - 1990:4.

In the first stage of analysis, the authors begin by using a three variable system consisting of nominal income, a fiscal variable and a money variable to estimate a VAR. Using the results, they use Granger causality tests of the null hypothesis that all coefficients on the lagged growth rates of money are equal to zero in the nominal income equation. For the first sample period the null hypothesis is rejected for the monetary base, M1, and M2. When the sample is expanded to 1990, the null cannot be rejected for the monetary base. For the third sample, the null can only be rejected for M1. Removing the fiscal variable yields similar results. What's more, when authors expand the data set to include the price index as well, the null hypothesis is rejected for M1 and M2 in the first two samples, but cannot be rejected for any money aggregate in the final sample. They argue that the experience of the 1980s seems to have altered previous empirical relationships between money and nominal income.

The second stage continues this analysis by examining the stability of money demand for the same three samples above with the variables expressed in levels rather than differences.[5] For example, a typical long run money demand function is given by:

$$m_t - p_t = \gamma_0 + \gamma_y y_t + \gamma_r r_t + e_t \tag{3}$$

where m is the money supply, p is the price level, y is a scale variable of real economic activity, r is the interest rate, and the variables are expressed in logarithms. For money demand to be considered stable in the long run, any deviations in money demand must be temporary.

The problem in estimating equation (3) is that all variables follow non-stationary I(1) processes and as a result have no tendency to return to a long run level. Nevertheless, it remains possible to examine the stability of money demand. For example, if deviations from equation (3) are temporary, e_t should be stationary. This will be the case if the I(1) variables are cointegrated, or share a common trend.

Friedman and Kuttner test the null hypothesis of no cointegration using the Johansen maximal eigenvalue likelihood ratio statistic using the unrestricted model above and by imposing two separate restrictions on equation (3); namely, a unitary income elasticity ($\beta = 1$) and an exclusion of the interest rate ($\gamma = 0$).[6] For the unrestricted case, they reject the null of no cointegration for the monetary base, M1, and M2 in the first sample. The null hypothesis is rejected only for M2 in the second sample and cannot be rejected for any

[5] It is important to consider the level specification because of the potential for lost information when the data is first-differenced. This point was recognized by Friedman and Kuttner.

[6] A restriction of $\alpha = 0$ is imposed on all models as the authors do not make explicit use of a constant term in the paper.

measure of money in the final sample. Similar results hold for the restricted models. As a result, Friedman and Kuttner (1992: 490) conclude that, "whatever the situation may have been before the 1980's, it is not longer possible to discern from the data a stable long-run relationship between income and the monetary base, M1, or credit, either with or without allowance for the effect of interest rates, and the evidence of such stability in the case of M2 strictly depends on the inclusion of data from the 1960's."

More recently, Estrella and Mishkin (1997) have used an approach similar to that of Friedman and Kuttner to examine the role of money growth in determining inflation and nominal income growth. Specifically, they estimate a VAR model that includes nominal income growth, inflation, and either the monetary base of M2 as the preferred measure of money growth using monthly data.[7] The data set covers the period 1960:3 - 1995:12. Also, the model is estimated for a subsample for the period beginning in October 1979 that coincides with the appointment of Paul Volcker as chairman of the Federal Reserve and is an important break point for the analysis of simple sum money aggregates.

Using the estimates from the three variable VAR, the authors conduct Granger causality tests of the null hypothesis that the lags of a given variable are all equal to zero. In estimation over the entire sample using the monetary base, the null hypothesis is rejected for the influence of lagged money growth on both nominal income growth and inflation. What's more, nominal income growth and inflation do not predict money growth. However, when the model is estimated in the subsample beginning in October 1979, the null hypothesis the coefficients on lagged money are equal to zero cannot be rejected in the nominal income or inflation equations. In fact, the null hypothesis is only rejected for the own lags of each variable in the subsample.

The results using M2 are not promising either. For the entire sample, lagged money does help to explain the growth in nominal income, but not inflation. Additionally, lagged values of inflation and nominal income growth are found to influence money growth. In the

[7] Inflation is thus defined as the change in the consumer price index. Nominal income growth is composed of the Commerce Department's index of coincident indicators and the consumer price index.

subsample, lagged values of money growth do not help to predict nominal income growth or inflation growth. Finally, lagged values of inflation help to predict the movements in money growth.

Overall, these results do not support the notion that money growth is a useful predictor of nominal income growth and inflation in the period since 1979. Nevertheless, it is possible that the poor performance of money as an information variable could be the result of countercyclical movements in money as a result of attempts to smooth fluctuations in inflation and nominal income growth. Estrella and Mishkin investigate this claim by measuring the size and significance of the sum of the coefficients on lagged nominal income growth and inflation in the money growth equation. The results show that the coefficient sum is often either not statistically significant or has the wrong sign. This is the case for both the monetary base and M2. As a result there is little reason to believe that the changes identified in the Granger causality tests are due to countercyclical movement in the money aggregates.

Whereas the literature discussed above focuses on the ability of money growth to predict nominal output growth and inflation, a second major empirical claim of the cashless approach is that money does not provide any additional information to explain fluctuations in the output gap. Empirically, this hypothesis can be tested by estimating the IS equation outlined above. Such empirical analysis has been conducted by Rudebusch and Svensson (2002) using a backward-looking IS equation of the form:

$$y_t = \beta_1 y_{t-1} + \beta_2 y_{t-2} + \beta_3 (i_{t-1} - \pi_{t-1}) + \varepsilon_t \tag{4}$$

where y is the output gap defined as the percentage deviation of real output from the Congressional Budget Office's measure of potential, i is the federal funds rate, and π is the average rate of inflation rate as measured by the GDP deflator.

The parameter estimates are obtained using a sample of quarterly data from 1961 - 1996. They find that the output gap has a strong autoregressive component and is negatively related to the lagged real interest. All three parameters are statistically significant and they report that these estimates are stable over time.[8] Notably missing from this analysis is

[8]The specific results are listed below in direct comparison to the empirical analysis in this paper.

money as the authors (ibid: 423) acknowledge that, "lags of nominal money (in levels or growth rates) were insignificant when added" to the IS equation above. These results are consistent with a complementary VAR approach used by Gerlach and Smets (1995) that suggests that money aggregates fail to provide additional information when added to an endogenous vector of output, inflation, and the interest rate.

The results outlined above cast serious doubt on the ability of money aggregates to explain economic activity. The evidence suggests that money demand is unstable and monetary aggregates are unable to explain movements in prices and nominal income. What's more, the lack of an identified role for money in explaining deviations in the output gap imply that movements in money aggregates are not useful as an information variable.

Nevertheless, there are reasons to be skeptical of this analysis. For example, recent estimates by Nelson (2002) show that lags of the real monetary base do have a positive and statistically significant effect on the output gap when added to Rudebusch and Svensson's IS equation. Similarly, Hafer, Haslag, and Jones (2007) show real M2 has a positive and statistically significant impact on the output gap independent of the real interest rate even in the subsample that begins in the 1980s. What's more, Hoffman, Rasche, and Tieslau (1995) show that the demand for real M1 is stable when a unitary income elasticity is imposed on the data. Also, Anderson and Rasche (2001) find that the demand for the real monetary base is stable using annual data from 1919 - 1999.

This paper similarly argues that the results outlined above should be met with skepticism. However, contrary to others, this paper suggests that the failure to identify stability in the demand for money and a role for money in business cycles is the result of the mismeasurement of the money aggregates. The idea of mismeasurement and its implications are the subject of the remainder of the paper.

3 Mismeasurement?

The vast majority of the empirical literature that estimates money demand functions and the effects of money on real economic activity employs simple sum money aggregates in which different monetary components are added together with equal weights. This procedure has long been considered inadequate for measuring money.[9] For example, in assessing different measures of money included in their *Monetary Statistics of the United States*, Friedman and Schwartz (1970: 151) noted that it would be more appropriate for the components of money aggregates to be assigned a weight based on their degree of "moneyness."

The reason that the weights of each asset are important is because the simple sum money aggregates imply that each asset is a perfect substitute for all other assets in the index. This is problematic because it is contrary to empirical evidence (Cf. Barnett, Fisher, Serletis, 1992; Serletis, 2001) and as a result simple sum aggregates fail to capture pure substitution effects across assets. The failure of simple sum aggregates to capture these substitution effects is important as it necessarily implies that there has been some change in the subutility function pertaining to monetary services and thus, potentially, the observed instability of money demand discussed above.

An alternative to the simple sum aggregates is the monetary services index derived by Barnett (1980) in which monetary assets are weighted by their expenditure shares.[10] Formally, this can be expressed as follows:

$$dlogM_t = \sum_{i=1}^{n} \bar{w}_{it} dlogx_{it}$$

where $\bar{w}_{it}$ is the expenditure share averaged over the two periods and x_{it} is the quantity of component i at time t. The numerator of the expenditure share, w_{it}, is the product of

[9] The earliest critic of simple arithmetic index numbers is likely Irving Fisher (1922).

[10] These aggregates have often been called "Divisia" aggregates in the literature because they are constructed using the Divisia method of aggregation. The term monetary services index is the name chosen by the St. Louis Federal Reserve in the official publication of the data (Anderson, Jones, and Nesmith, 1997a; 1997b). This name is meant to reflect the fact that these aggregates measure a flow of services from a class of assets rather than a stock of assets.

the user cost of the particular asset and dollar quantity of that asset. The denominator is the summation of these products over all assets in the index. Here the user cost of asset is derived from Barnett (1978) as:

$$u_{it} = \frac{(R_t - r_{it})}{(1 + R_t)} P_t$$

where u_{it} is the nominal user cost of asset i at time t, R is the benchmark rate of return, r_i is the return on asset i, and P is the price level.[11]

The derivation of the monetary services index (henceforth MSI) is important for two reasons. First, these aggregates are derived from explicit microfoundations and are consistent with aggregation and index number theory. Second, the MSI aggregates are capable of adapting to financial innovation both through the introduction of new money assets or a change in the interest-bearing properties of a given asset. Simple sum indexes do not satisfy either criteria.

Despite the clear theoretical superiority of the weighted money aggregates, it is not clear a priori that this necessarily implies a corresponding quantitative difference with simple sum aggregates. In order to facilitate such a comparison, the differences between MSI M1, MSI M2, MSI MZM, and the simple sum counterparts are plotted in Figures 1 - 3.[12] It is important to note that differences in the growth rates are most notable in the 1980s, the decade in which money supposedly became less useful as both an intermediate target and an information variable.[13] In addition, the growth of simple sum M2 is greater than the MSI

[11]Donovan (1978) argues that the user cost concept is more appropriate for determining the price of money than the traditional form of assigning a price of unity.

[12]Throughout the paper, MSI M1, MSI M2, and MSI MZM are used as money aggregates. In choosing these aggregates, this paper explicitly ignores a second problem with monetary aggregation, which is the composition of assets within the aggregate. In other words, assigning weights to assets within an aggregate is not sufficient for designing a valid aggregate. The emphasis in this paper is in providing analysis with aggregates in which the composition is well-known.

[13]This point is potentially of importance. In the examination of simple sum aggregates, some such as Carlson, Hoffman, Keen, and Rasche (2000) have argued that changes in stability of money demand driven by structural shifts should be adjusted accordingly. Others, such as Woodford (1998), have argued that such shifts are what make money demand unstable and unusable.

counterpart for most of sample.

These differences in growth rates are, in fact, quantitatively important. For example, Belongia (2005) finds that during the 1960s and 1970s, the differences between the growth rates of simple sum M1 have a unit root. Thus, following the simple sum aggregate could potentially result in vastly divergent predictions than the MSI counterpart.

What's more, these differences seem to be most important during the time in which money is thought to have lost its predictive ability. Barnett (1997) highlights the so-called monetarist experiment of 1979 - 1982 as well as the remainder of the early 1980s as two such instances. Specifically, Barnett highlights the fact that simple sum M2 and M3 grew at an average rate of 9.3% and 10%, respectively, during the monetarist experiment whereas the MSI counterparts grew at 4.5% and 4.8%. These average rates came on the heels of double-digit growth rates for both the simple sum and MSI aggregates. Thus, using the MSI aggregates, it is much easier to understand why a severe recession rather than a mild disinflation resulted from the contractionary policy.

The second departure can explain why dire monetarist predictions of rising inflation in the subsequent period were incorrect. The sudden increases in money growth in the early 1980s was the result of new financial assets, such as NOW accounts and money market deposit accounts. As noted by Barnett (1997), the differences in the growth rate of the monetary aggregates can be explained by the way in which new assets are introduced. New assets are simply added to simple sum aggregates. In contrast, new assets are introduced to MSI aggregates using the appropriate weight. Thus, given that the interest rates on these new assets were relatively high, the weight was correspondingly low thereby making for a smooth introduction to the MSI aggregate. Note that this change is highlighted by the largest spike in the growth rate differences shown in Figures 1 - 3.[14] It is also important to highlight the fact that among the aggregates the difference is largest for MZM, which includes both new

[14]Note that even if one recognized that this was the reason for the spike in the growth rate in simple sum aggregates, it would still be difficult to assess how much of the change was the result of the introduction of new assets and how much was the result of monetary policy.

assets.[15]

In addition to the differences highlighted by casual inference, recent empirical evidence does suggest that the way in which money is measured has important implications for one's results. For example, Belongia (1996) re-examines five puzzling results from the monetary literature by utilizing MSI aggregates rather than their simple sum counterparts. He finds that four of the five puzzling results exist only when simple sum aggregates are used. The results are mixed for the fifth result. In addition, the work collected in Belongia and Binner (2000) shows that MSI aggregates outperform their simple sum counterparts for most countries.

Given these results, it is important to re-examine the empirical evidence using the MSI measures of money rather than the simple sum indexes, which are at best theoretically flawed and at worst empirically misleading. This re-examination is the subject of the next section.

4 Empirical Evidence

Overall, the results presented in section 2 raise doubts about the stability of money demand as well as any information role that monetary aggregates might possess. This section re-examines those results by using the monetary services indexes outlined above to measure money while employing the same methods as earlier research. The first subsection analyzes the stability of the demand for real money balances. The second and third subsections consider the role of money as an information variable.

4.1 Money Demand Stability

Recall the long run money demand equation outlined in section 2 above:

$$m_t - p_t = \gamma_0 + \gamma_y y_t + \gamma r_t + e_t$$

where m is the money supply, p is the price level, y is a scale variable of real economic activity, r is a price variable usually measured by an interest rate, and the variables are

[15] M2 also includes both assets, but also includes time deposits that are not included in MZM.

expressed in logarithms.

As previously mentioned, given that each of these variables are non-stationary, there must exist a linear combination of the variables that is stationary for money demand to be considered stable in the long run. In other words, money demand stability requires that deviations from equilibrium, as expressed by the money demand function, are temporary.

In order to determine whether there exists a stable money demand function, it is therefore useful to employ the cointegrated VAR approach. The use of this approach is important because it not only allows one to test for cointegration, but also to estimate the parameters of the money demand function. Formally, the p-dimensional cointegrated VAR model is given by:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_k \Delta x_{t-k} + \Pi x_{t-1} + \varepsilon_t$$

where x_t is a vector of non-stationary I(1) endogenous variables, and Γ_i, $i = 1, \ldots, k$, and Π are (p x p) parameter matrices. The matrix, Π, captures the long-run properties of the model. Within this context, the cointegration hypothesis is expressed as a reduced rank restriction on Π, which can be written as the product of two matrices:

$$\Pi = \alpha\beta'$$

where α and β are (p x r), $r \leq p$, matrices of short-run adjustment coefficients and long run equilibrium coefficients, respectively, and $rank(\Pi) = r$. As a result, the cointegrated VAR can be written:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_k \Delta x_{t-k} + \alpha\beta' x_{t-1} + \varepsilon_t \tag{5}$$

where $\beta' x_t$ is an (r x 1) vector of cointegrating relationships. It follows that the rank of Π is equal to the number of cointegrating vectors. The existence of a stable long run money demand function is therefore consistent with a cointegrating vector:

$$\beta' x_t = (m_t - p_t) - \gamma_0 - \gamma_y y_t - \gamma_r r_t = 0$$

in which the parameters γ_0, γ_y, and γ_r are constant across subsamples.

Before estimating the model it is important to determine how the variables are measured. Real money balances are measured by MSI M1, MSI M2, and MSI MZM adjusted for the

GDP deflator. Typically, the scale variable is measured by some measure of real economic activity such as real gross domestic product. However, given that money demand is derived from consumer choice theory, real GDP is not likely to be the proper measure of income even for a representative agent. As a result, the scale variable used in this paper is the real final sales of domestic production.[16]

The own price of money is generally proxied by the use of the opportunity cost of holding money, which is often measured by a short term interest rate.[17] Ultimately, however, the use of the interest rate as the price of money is incorrect. As Belongia (2006: 240) notes, this "confuses the concepts of 'credit' and 'money'." The appropriate measure of the own price of money when using a monetary services index is straightforward as it is given by the price dual of the monetary services index.[18,19]

With the variables now properly defined, the cointegrated VAR model is estimated over the sample that spans 1960 - 2005 for each measure of money. The lag length is determined by Akaike information criteria. The stability of money demand is determined first by using Johansen's trace statistic to identify the rank of Π and thus the number of cointegrating vectors. Next, given that the ability to identify cointegration is dependent on the sample in the earlier work cited above, the model is estimated recursively by choosing a subsample, $1, \ldots, T_1$, where $T_1 < T$, and then extends the end point of the subsample until the entire sample is estimated. Using the results from the recursive estimation, one can test whether the

[16]The results are not sensitive to this specification.

[17]Poole (1988) suggests that a long term interest rate should be used. Hoffman, Rasche, and Tieslau (1995) note that it is of little consequence in a cointegrating VAR model as such interest rates are typically cointegrated and thus adding an additional interest rate would simply result in an additional cointegrating vector.

[18]A price index number is the dual of the quantity index when the product of the quantity index and the price dual equal the total expenditure on the assets in the quantity aggregate. This can be thought of as the price of one unit of monetary services. For more, see Anderson, Jones, and Nesmith (1997b).

[19]It remains possible that an interest rate can serve as an opportunity cost variable that shifts the demand curve. However, using the yield on the 90-day Treasury bill, the hypothesis that the interest rate could be excluded from the money demand function as tested by a restriction on the cointegrating vector could not be rejected. As such, the interest rate is excluded from the results below.

parameters and the number of cointegrating relationships are constant across samples. This enables a comparison with the results of Friedman and Kuttner (1992) without specifying one particular subsample.

The first stage of the analysis is to identify the rank of Π, and thus the number of cointegrating vectors, over the entire sample. The corresponding Johansen trace statistics are shown in Table 1. For each definition of money, the null hypothesis of no cointegration is rejected. What's more, the null hypothesis that $r \leq 1$ cannot be rejected for any measure of money. The existence of one cointegrating vector suggests that money demand is stable for the entire sample.

With the rank of Π identified, the model is now re-estimated by imposing the restriction, $rank(\Pi) = 1$. With the coefficient on real money balances normalized to unity, the corresponding coefficient estimates of the cointegrating vector are shown in Table 2. The coefficients are generally thought to follow the conditions, $\gamma_y > 0$ and $-1 < \gamma_r < 0$. As one can see from Table 2, these conditions are satisfied for each measure of money.

It is now important to consider the stability of the coefficient estimates as well as the number of cointegrating relationships over time. Parameter stability is examined using the max test of a constant β defined in Juselius (2006), which tests the null hypothesis that $\hat{\beta}_{t_1} = \hat{\beta}_T$, where $t_1 = T_1, \ldots, T$ is the endpoint of the recursive subsample. The stability of cointegrating relationships are measured using recursively calculated trace statistics:

$$\tau(j) = \Big\{ -t_1 \sum_{i=1}^{j} ln(1 - \hat{\lambda}_i) \Big\} \qquad j = 1, \ldots p; t_1 = T_1, \ldots, T$$

where λ_i, $i = 1, \ldots, p$, are the eigenvalues of Π.

Each of these tests is calculated using the X-form, the complete model shown in equation (5) above, and the R-form, in which the short-run dynamics, $\Gamma_i, i = 1, \ldots, k$, of the standard VAR model have been concentrated out. The importance of estimating the test statistics for each model is because, in the R-form, the effects of non-constant parameters have been averaged out and thus might more accurately reflect whether the parameters are constant in the long-term structure of the model.[20] What's more, to facilitate easier analysis the test

[20]The derivation of the R-form is shown in the Appendix and in Juselius (2006) and Hansen and Johansen

statistics are presented in graphical form for all recursive subsamples and are divided by the 95% quartile such that whenever the null hypothesis is rejected, it will be illustrated graphically as a value greater than unity.

Figures 4, 5, and 6 plot the recursive trace statistics for MSI M1, MSI M2, and MSI MZM, respectively. For each measure of money, the null hypothesis of $r = 0$ is rejected for every period in the recursive sample when calculated based on the X-form of the model. What's more, the null hypothesis of $r \leq 1$ cannot be rejected for MSI M1 for every period in the recursive sample. This hypothesis can be rejected for a brief period in the 1970s for MSI M2 and is nearly rejected for MSI MZM during the same period. However, this hypothesis cannot be rejected for the R-form and therefore suggests that the rejection is driven by the short run rather than the long run dynamics of the model. Perhaps more interesting is that, for the R-form of the model, the null hypothesis of $r = 0$ cannot be rejected until the early 1980s for MSI M2 and MSI MZM and the early 1990s for MSI M1. Thus, the trace statistics for the X-form largely suggest that there exists one cointegrating relationship over the recursive sample thereby providing evidence of money demand stability. In contrast, the R-form of the model suggests, at least at first glance, that the existence of a stable money demand function might be dependent on the sample. However, these latter results might merely reflect the low power of the trace tests and as a result cause one to erroneously reject a stationary linear relationship. Upon casual inspection, this would appear to be the case as one needs a large sample to correctly reject the hypothesis of a unit root and the trace statistic appears to be monotonically increasing as the sample size grows.

An alternative explanation of the differing results of the X-form and the R-form of the model is that there exists long run instability in the model.[21] Luckily there exist alternative tests to examine whether the failure of the trace statistics to reject the null hypothesis are the result of unstable relationships or sample size.

(1999).

[21]Note that the R-form of the model captures the long-run by concentrating out the short-run dynamics of the standard VAR. Thus the failure to reject the null of $r = 0$ for the R-form and not the X-form must result from the long-run dynamics of the model.

The first procedure is the eigenvalue fluctuations test that is designed to measure the constancy of the individual eigenvalues of Π across recursive samples. Note that the recursive trace statistic outlined above is a function of the sample size and the eigenvalues of Π. What's more, these eigenvalues can be expressed as quadratic functions of the adjustment and cointegrating matrices. It therefore follows that if there are changes in the trace statistic across recursive samples, this is either the result of the sample size or long run instability of the model. The eigenvalue fluctuations test examines whether the eigenvalues are constant across recursive samples in an attempt to identify the nature of the problem. If the eigenvalues are constant, changes in the trace statistic must be the result of sample size. If, on the other hand, the eigenvalues are not constant, then changes in the trace statistic must be the result of some sort of long run instability in the model, or in the case of this particular model, the instability of money demand.

The test statistic for the fluctuations test is expressed as:

$$\tau_i(t_1) = \frac{t_1}{T}\sqrt{T}\Sigma_{ii}^{-1/2}(\hat{\lambda}_{i,t_1} - \hat{\lambda}_{i,T})$$

where Σ_{ii} is the variance of λ_i defined by Hansen and Johansen (1999), λ_{i,t_1} is the eigenvalue i, $i = 1, \ldots, p$, of the recursive sample t_1, and $\lambda_{i,T}$ is the eigenvalue i of the complete sample. For the model estimated here, $rank(\Pi) = 1$ and thus there is one eigenvalue to examine for each money demand equation. These test statistics have been divided by the 5% critical value and are plotted in Figures 7 - 9. As shown, the null hypothesis of the constancy of the eigenvalues cannot be rejected across recursive samples for any measure of money. These results provide further evidence that the failure to reject the hypothesis of no cointegration in the early recursive samples is the result of sample size rather than instability.

The second examination procedure is the max test for β constancy, which tests the null hypothesis that the estimates of β for each of the recursive samples is equal to that of the entire sample. This test is important for two reasons. First, even if one can reject the null hypothesis of no cointegration, money demand can still be considered unstable if the corresponding parameters are not constant across different subsamples. Second, if one cannot reject the null hypothesis of constancy, it would lend support to the claim that the failure to identify a cointegrating vector in the R-form of the model is the result of the weak power of the test rather than money demand instability.

The results of the max test for β constancy are illustrated in Figures 10, 11, and 12 for MSI M1, MSI M2, and MSI MZM, respectively. For every measure of money, one cannot reject the null hypothesis of a constant β for both the X-form and the R-form of the model over all recursive samples. Given that the model is estimated after imposing $rank(\Pi) = 1$, this implies that the parameters of the money demand function are stable for each measure of money over all of the samples estimated. In addition, this identification of long run stability in the parameters of β strongly indicate that the failure to identify a cointegrating vector for early recursive samples is the result of the small sample size and the low power of the test rather than some form of long run instability in the model.

These results are important as they contrast significantly with those of Friedman and Kuttner (1992). Whereas their earlier work suggested that money demand was not stable since the 1980s, the evidence presented above suggests that the stability of money demand is an empirical reality when a monetary services index is used as the monetary aggregate. For each measure of the monetary services index, there does exist a single cointegrating vector with reasonable and stable parameter values for a money demand function across recursive samples. What's more, the only instances in which the null of no cointegration cannot be rejected are instances in which the sample size is too small as evident in the eigenvalue fluctuations test above. These results therefore provide credence to the hypothesis that empirical failures relating to money demand are an issue of mismeasurement.

4.2 Money, Income, and Prices

For money to serve as an information variable it is important to determine whether money growth can help predict inflation and nominal income growth. This approach is taken by Friedman and Kuttner (1992) as well as Estrella and Mishkin (1997). This section adopts the three variable system of Estrella and Mishkin, estimates the corresponding VARs, and conducts Granger causality tests of the null hypothesis that lags of a given variable do not effect the particular variable in question. Nominal income growth is measured by nominal GDP growth, inflation by the change in the GDP deflator, and money growth by the monetary services indexes. To facilitate comparison with Estrella and Mishkin, the model is estimated over two samples, with the break point occurring in October 1979.[22]

Tables 3 - 5 presents the p-values of the Granger causality tests for the three variable system for the pre-1979 sample. The results are printed such that the null hypothesis is that the column variable does not have an effect on the row variable. These results suggest that money, as measured by MSI M2 and MSI MZM, does help to predict nominal GDP growth. What's more, the lags of all measures of money help to predict the inflation rate.

The p-values for the Granger causality tests for the post-1979 era are shown in Tables 6 - 8. These results are different both from the pre-1979 results and those shown in Estrella and Mishkin for the same period. In contrast to the earlier sample and consistent with the work highlighted above, each measure of the monetary aggregate cannot predict fluctuations in nominal GDP growth. Conflicting with the work of Estrella and Mishkin, however, is that all of the money aggregates are useful for predicting inflation.

While the ability of the growth rates of the monetary services indexes to predict inflation is a notable improvement over the results shown in Estrella and Mishkin (1997), it remains somewhat puzzling that the same aggregates cannot predict nominal income growth in this latter period. One potential reason for this failure could be the result of the fact that movements in the money supply are reflecting an increased responsiveness of monetary policy to fluctuations in nominal income growth post-1979.[23] For example, if the central bank is responding to both changes in inflation and the output gap, as is a generally accepted proposition for the United States, this will likely be reflected in the bank's responsiveness to nominal income growth. Upon casual inspection, this is potentially the case as nominal income growth does help predict the growth rates of two of the three money aggregates in the VAR model.

[22]The lag length for the VARs are determined by lag length specification tests.

[23]Hendrickson (2010) shows this to be the case, but uses the federal funds rate as the measure of monetary policy rather than a measure of the money supply.

This hypothesis is tested directly by estimating a central bank reaction function in which the bank's intermediate target is a money aggregate and its target is nominal income growth. Formally, this can be tested by estimating the following regression:

$$\Delta M_t = \alpha + \beta \Delta x_t + e_t$$

where ΔM_t is the growth rate of the money supply, Δx_t is nominal income growth, and α is a constant.[24] For this regression, nominal income growth is measured by the Federal Reserve's Greenbook forecast of nominal income growth.[25] The use of the forecast of nominal income growth is important because it eliminates the possibility of capturing reverse causation while also using data that was available to policymakers at the time.

The results of this regression are shown in Table 9.[26] The results indicate that the Federal Reserve forecast of nominal income growth did have a negative and statistically significant effect on the growth rate of MSI M2 and MSI MZM in the post-1979 era. This therefore lends credence to the claim that the failure of monetary aggregates to predict nominal income is the result of the fact that changes in nominal income have feedback effects on money aggregates through monetary policy.

A second way to examine whether monetary policy can explain why money growth cannot predict nominal income growth is to consider a case in which money aggregates are truly exogenous. For example, Rowe and Rodriguez (2007) find that changes in the growth rate of U.S. simple sum monetary aggregates do not Granger cause fluctuations in the growth rate of U.S. real GDP. However, changes in the growth rate of U.S. simple sum monetary aggregates do explain fluctuations in the real GDP of Hong Kong, whose currency is pegged to the U.S.

[24] Formally, this model could be re-written as:

$$\Delta M_t = \delta_0 + \delta_1(\Delta \bar{x} - \Delta x_t) + e_t$$

where $\bar{x}$ is the nominal income growth target. Thus, in the regression above, $\alpha = \delta_0 + \delta_1 \Delta \bar{x}$ and $\beta = -\delta_1$.

[25] This data is readily available through the Philadelphia Federal Reserve. The sample estimated is from 1979:4 - 2003:4 as that is the latest data available.

[26] The t-statistics correspond to Newey-West standard errors due to the evidence of serial correlation in a standard, ordinary least squares regression.

dollar. Intuitively, this is the case because U.S. monetary policy is reacts to fluctuations in measures of real activity in the U.S., but not Hong Kong. As a result, changes in the money supply in the U.S. should be exogenous to Hong Kong.

Thus, as a further method of comparison, the three variable system outlined above is re-estimated using nominal income growth and inflation from Hong Kong and the MSI aggregates from the U.S. The sample period runs from 1983:1 - 1997:2.[27] The results are shown in Tables 10 - 12. Based on the results, the growth of monetary aggregates can be considered exogenous for MSI M1 and MSI M2 as these are not predicted by nominal income growth or inflation. In addition, the same monetary aggregates do help predict Hong Kong's nominal income growth and inflation rate. Again, this tends to lend credence to the view that the failure to identify a role for monetary aggregates in predicting nominal income growth is the result of monetary policy.

Overall, the results from the three variable VAR using the monetary services index as the method of aggregation for the money supply indicate that money growth is an important predictor of inflation for all measures of money and all sample periods. This is important because previous research that relies on simple sum monetary aggregates find no such relation in the post-1979 era. The results with regards to nominal income, however, are somewhat mixed. In the pre-1979 era, MSI M2 and MSI MZM do help to predict nominal income growth. However, consistent with earlier research, this relationship does not hold in the post-1979 era. Nevertheless, this failure in the latter period is likely the result of countercyclical monetary policy. As evidence in support of this claim MSI M2 and MSI MZM demonstrate a statistically significant response to the Federal Reserve's forecast of nominal GDP growth. What's more, the U.S. MSI aggregates do help to predict nominal GDP growth and inflation in Hong Kong, a country whose currency is pegged to the U.S. dollar. Ultimately, the results in this subsection cast doubt on earlier research that finds money curiously unable to

[27]The sample is chosen as representing the period from the first peg of the Hong Kong dollar to the U.S. dollar until the handing over of Hong Kong to the Chinese and the re-pegging that occurred shortly thereafter.

predict inflation. In addition, and again contrary to earlier research, the inability of money aggregates to explain nominal income growth is shown to be the result of countercyclical monetary policy.

4.3 Money and the Output Gap

The final method of empirical analysis is an examination of the role of real money balances in predicting movements in the output gap. In general, this analysis has been carried out by estimating a backward-looking IS equation. As previously mentioned, Rudebusch and Svensson (2002) estimate the particular specification given in equation (4) using quarterly data from 1961 - 1996. They find little evidence of a role for money. This paper extends their analysis in two ways. First, the sample is extended through 2005:4. Second, the model is expanded to include a one period lag of the quarterly growth rate of real money balances as measured by the monetary services indexes.[28]

In addition to the estimates over the entire sample, the IS equations are also estimated for the subsample, 1979:4 - 2005:4. This date is chosen because it is consistent with the break point identified by Friedman and Kuttner (1992) and Estrella and Mishkin (1997) in their work described above. What's more, this date marks the change in monetary policy beginning with the appointment of Paul Volcker that has been documented by Taylor (1999), Clarida, et. al (2000), and Hendrickson (2010) as well as the beginning of financial deregulation and innovation. Others, such as Bernanke and Mihov (1998), Leeper and Roush (2003) and Hafer, Haslag, and Jones (2007) use 1983 as the break point. This is justified by the fact that this date marked the shift from monetary targeting to interest rate targeting as well as shifts in the velocity of certain simple sum aggregates. Practically, this is not a particularly useful date, especially for the monetary services indexes, which do not experience structural

[28]Formally, the use of lagged money growth can be justified by the fact that under certain functional forms, one can derive an IS equation with real money balances. In addition, the use of real balances could be justified on the grounds that they are an informational variable that reflects the substitution effects resulting from a change in the nominal money supply that are not captured by changes in "the" interest rate.

shifts in velocity around 1983.

The model is estimated using OLS. The results are shown in Table 13. The first column are the results reported by Rudebusch and Svensson. The second and third columns re-estimate the model with for the extended sample and the subsample, respectively. In each case, the output gap is shown to be strongly autoregressive and the real interest rate is shown to have a negative and significant impact on the output gap. Although there is a slight reduction in the parameter on the real interest in the subsample beginning in October 1979, the parameters are relatively constant across samples.

Columns 4 - 9 extend the model to include lagged quarterly growth of real money balances as measured by MSI M1, MSI M2, and MSI MZM adjusted by the GDP deflator. For the entire sample real money balances exhibit a positive and significant effect on the output gap. The results from the subsample do not find a statistically significant effect of real MSI M1 on the output gap. However, real balances do have a positive and statistically significant effect on the output gap when measured by MSI M2 or MSI MZM. In addition, the coefficient on the real interest rate declines in the subsample when real money balances are included and is not statistically significant when MSI M2 or MSI MZM is included in the regression. Overall, these results not only do not support omitting real money balances from IS equations, but also suggests that the exclusion of money leaves the estimated IS equations misspecified.

5 Conclusion

There is an emerging consensus that monetary aggregates are not useful in monetary policy and business cycle analysis. This view has largely been justified by empirical work that shows that the demand for money is unstable and that money does not help to explain fluctuations in the output gap. Modern business cycle theorists have used these results to develop models that completely abstract from money. At the core of these models is the dynamic New Keynesian IS-LM-type model where the LM curve has been replaced by an interest rate rule followed by the central bank. Money is inconsequential to the model as it merely reflects movements in the interest rate. In other words, money is redundant.

One potential problem with the empirical results that justify these cashless models is that they rely on the use of simple sum monetary aggregates. Such aggregates are theoretically flawed in that they treat all components of a particular aggregate as perfect substitutes; a result inconsistent with empirical evidence. Thus, previous results that employ simple sum aggregates are potentially flawed by mismeasurement.

As a result, this paper re-examines the empirical findings of previous authors by using monetary services indexes rather than the simple sum counterparts. The advantage of using the monetary services index is that it is derived from microtheoretic foundations and is consistent with aggregation and index number theory. Using this alternative measure of money, this paper identifies a stable money demand function for each component class of monetary assets across recursive samples. In addition, it is demonstrated that real money balances not only have a positive and significant impact on the output gap, but that this effect is often larger in magnitude than that of the real interest rate. Overall, the results suggest that previous findings are likely the result of mismeasurement with regards to monetary aggregates.

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Appendix

The R-form of the cointegrated VAR can be derived as follows:[29]

Consider the following equation for estimation:

$$y = \alpha + \beta_1 x + \beta 2z + \varepsilon$$

Following the Frisch-Waugh theorem, one can derive the OLS estimate of β_2 from the following regression:

$$u_1 = \beta_2 u_2 + e$$

where $u_1 = y - \hat{b}_1 x$ and $u_2 = z - \hat{b}_2 x$

Recall the cointegrated VAR model:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_k \Delta x_{t-k-1} + \Pi x_{t-1} + \varepsilon_t$$

Define $\Delta x_t = Z_{0,t}$, $x_t^* = Z_{1,t}$, $(\Delta x'_{t-1}, \ldots, \Delta x'_{t-k-1}) = Z_{2,t}$, and $(\Gamma_1, \ldots, \Gamma_k) = \Gamma$ such that the model can be re-written:

$$Z_{0,t} = \alpha\beta' Z_{1,t} + \Gamma Z_{2,t} + \varepsilon_t$$

Applying the basic principle of the Frisch-Waugh theorem, the R-form of the model in which the short-run dynamics, Γ, have been concentrated out is thus given by:

$$R_{0,t} = \alpha\beta' R_{1,t} + R_{\varepsilon,t}$$

where

$$R_{0,t} = Z_{0t} - M_{02} M_{22}^{-1} Z_{2t}$$

$$R_{1,t} = Z_{1t} - M_{12} M_{22}^{-1} Z_{2t}$$

$$R_{\varepsilon,t} = \varepsilon_t - M_{\varepsilon 2} M_{22}^{-1} Z_{2t}$$

and

$$M_{ij} = \sum_{s=1}^{t} Z_{it} Z'_{jt}$$

$$M_{\varepsilon j} = \sum_{s=1}^{t} \varepsilon_t Z'_{ij}$$

[29]What follows is derived in Juselius (2006) and Hansen and Johansen (1999).

Figure 1: Differences in Growth Rates – M1

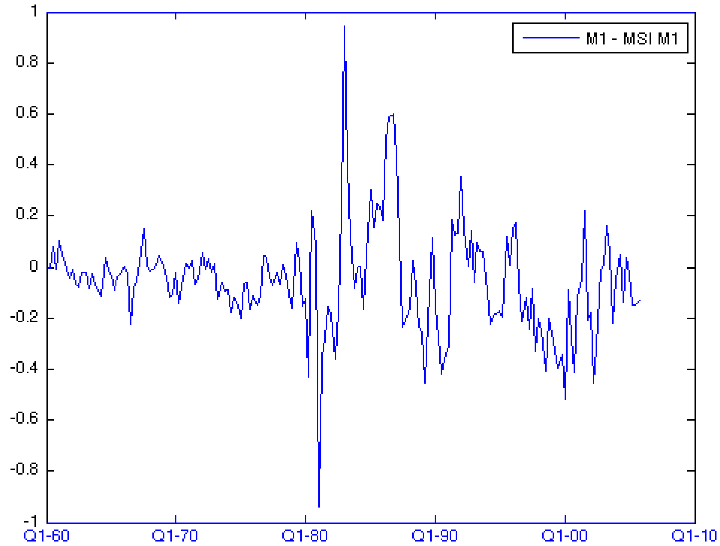

Figure 2: Differences in Growth Rates – M2

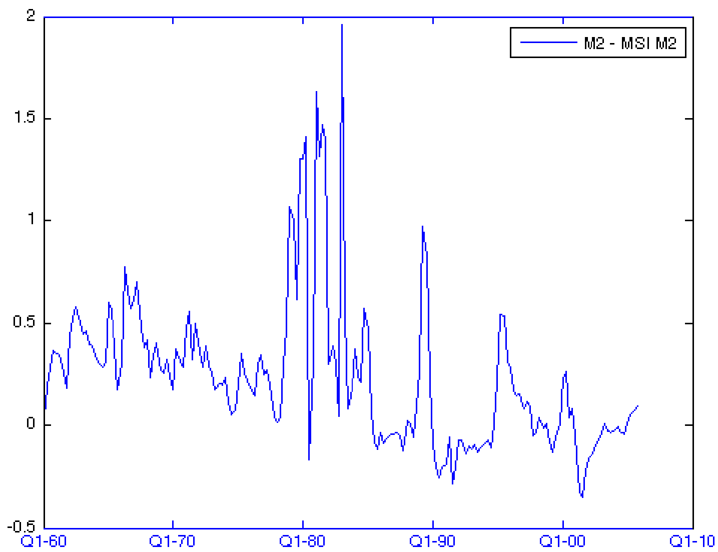

Figure 3: Differences in Growth Rates – MZM

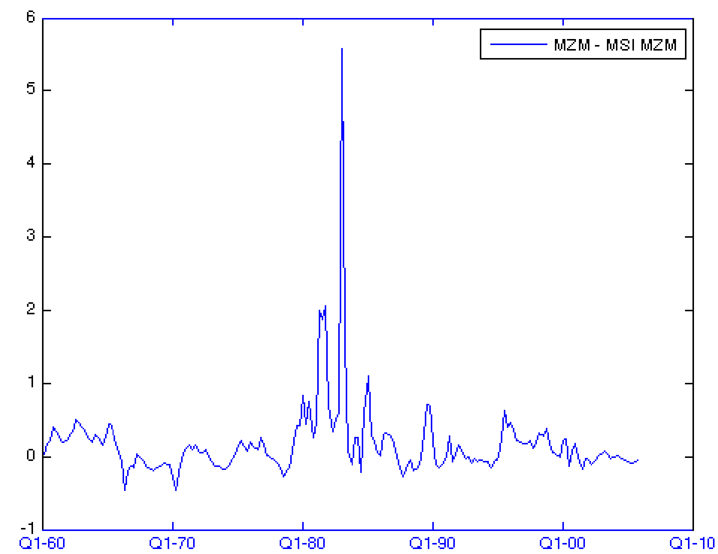

Table 1: Trace Statistics

Monetary Variable	r	Trace	P-Value
MSI M1	0	51.67	0.00
	1	11.95	0.46
	2	3.43	0.52
MSI M2	0	59.81	0.00
	1	9.91	0.65
	2	3.55	0.50
MSI MZM	0	59.31	0.00
	1	11.81	0.47
	2	2.68	0.65

Table 2: Cointegrated VAR Parameter Estimates

Monetary Variable	RFINSAL	DUAL	CONS
MSI M1	0.39	-0.72	1.18
MSI M2	0.64	-0.38	-1.53
MSI MZM	0.93	-0.97	-1.51

Figure 4: Recursive Trace Statistics – MSI M1

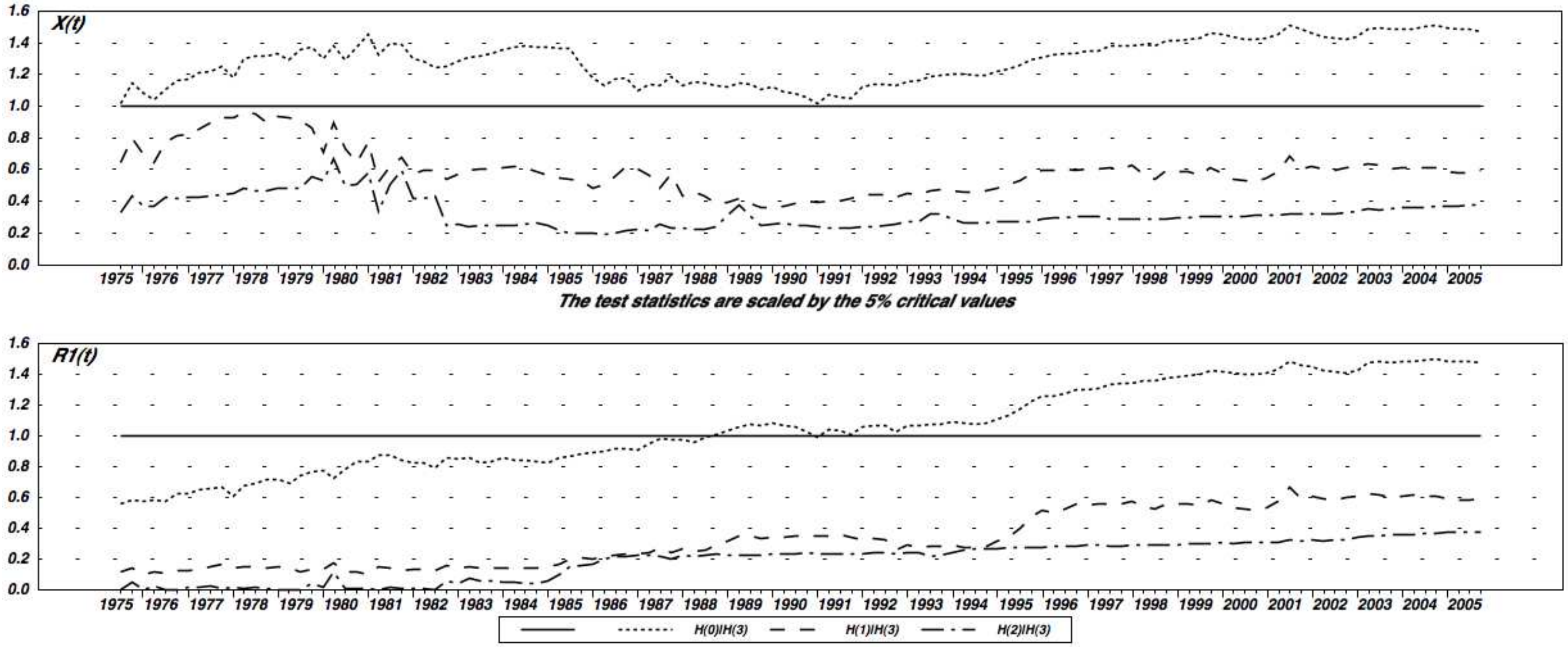

Figure 5: Recursive Trace Statistics – MSI M2

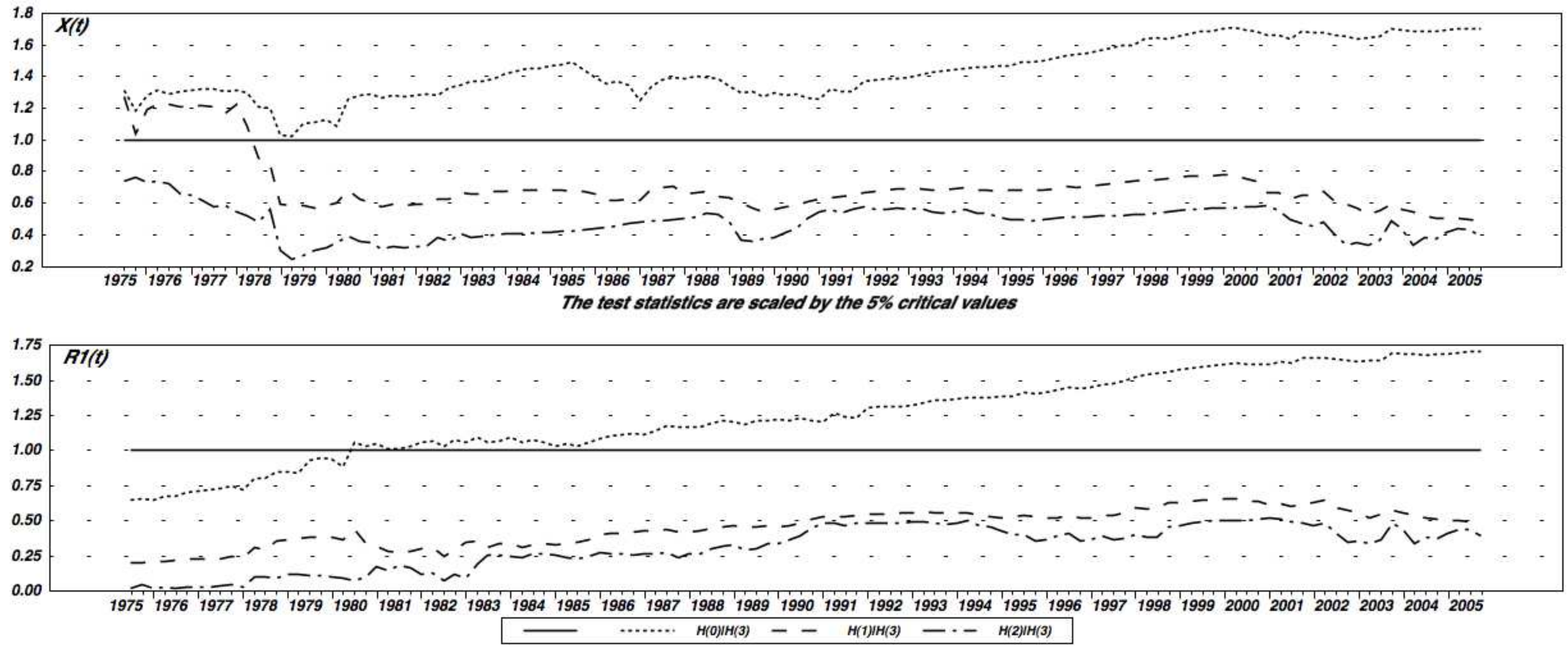

Figure 6: Recursive Trace Statistics – MSI MZM

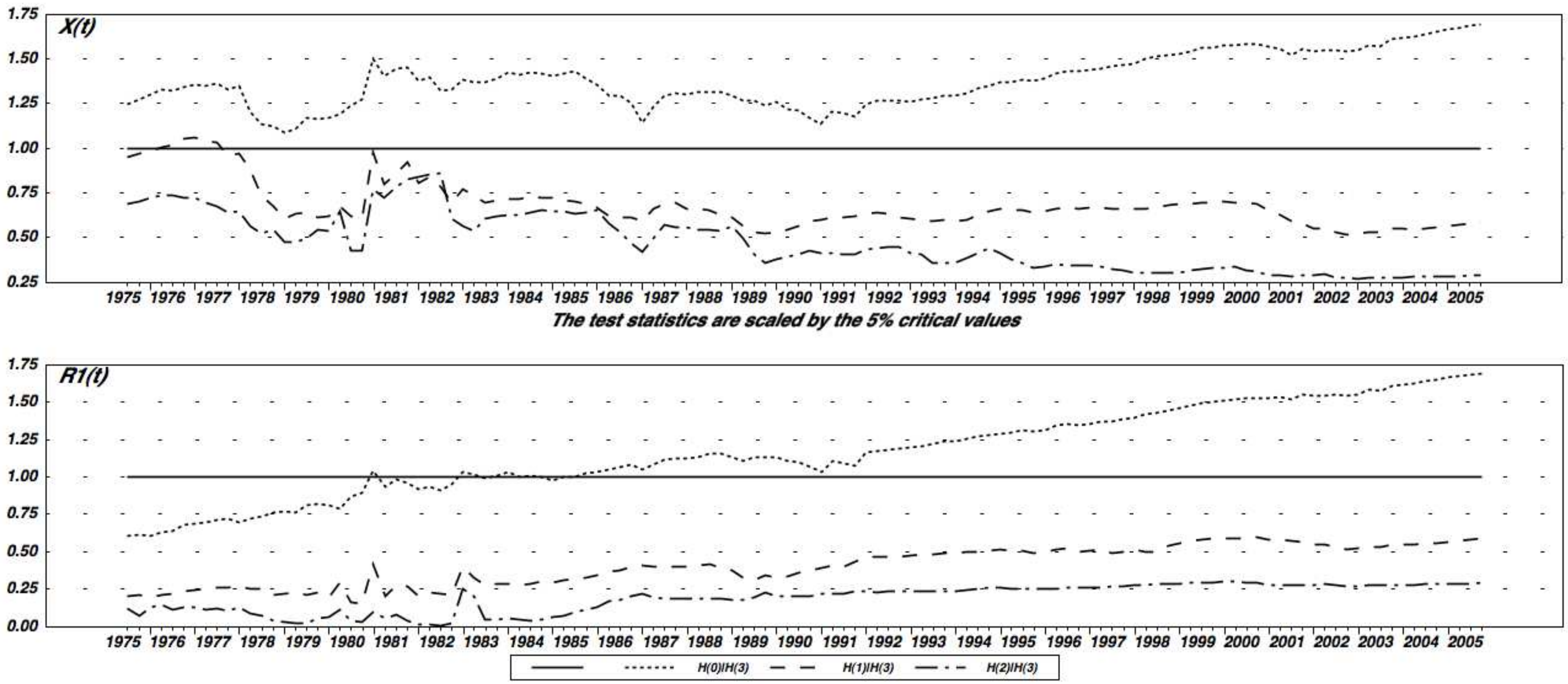

Figure 7: Fluctuations Test – MSI M1

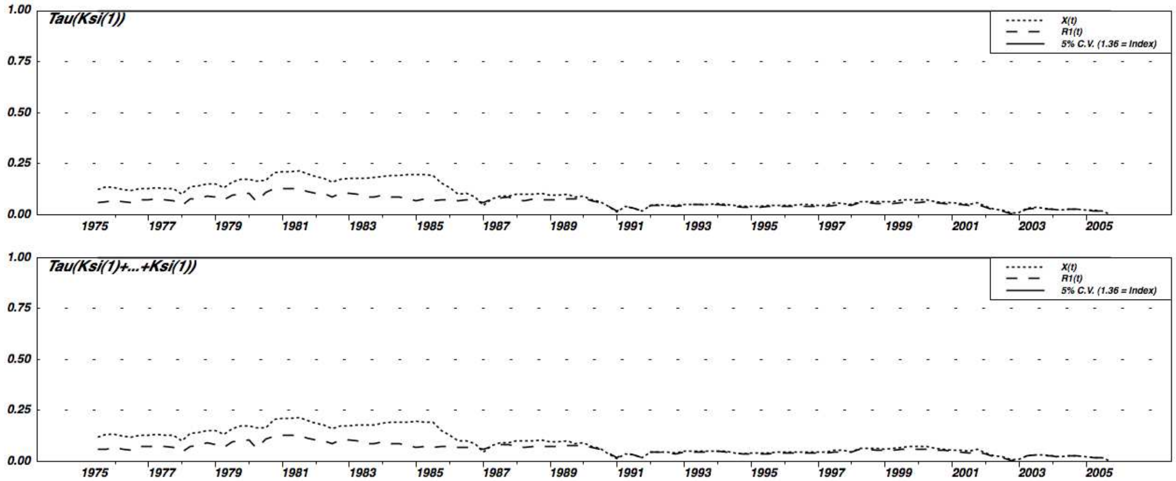

Figure 8: Fluctuations Test – MSI M2

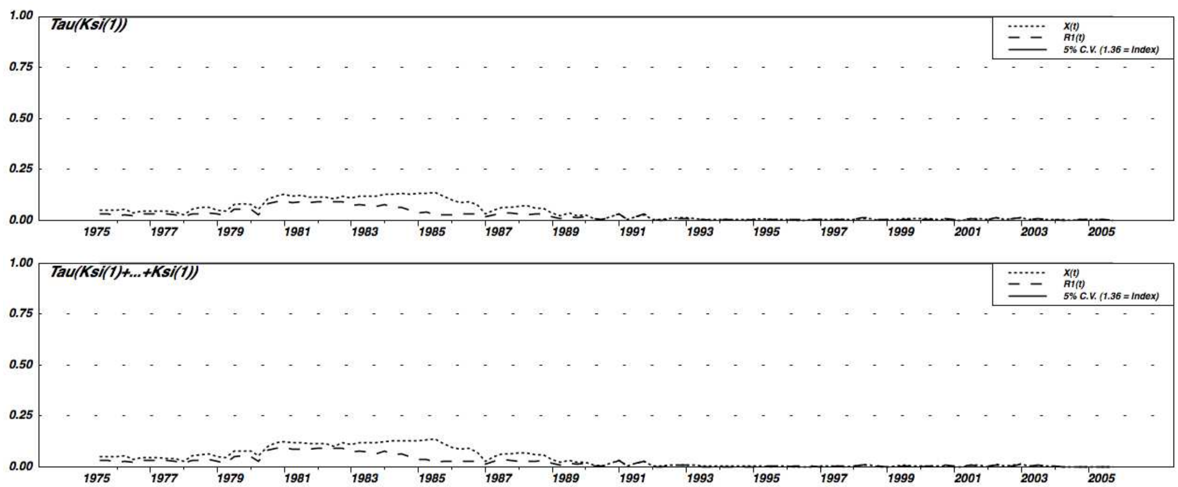

Figure 9: Fluctuations Test – MSI MZM

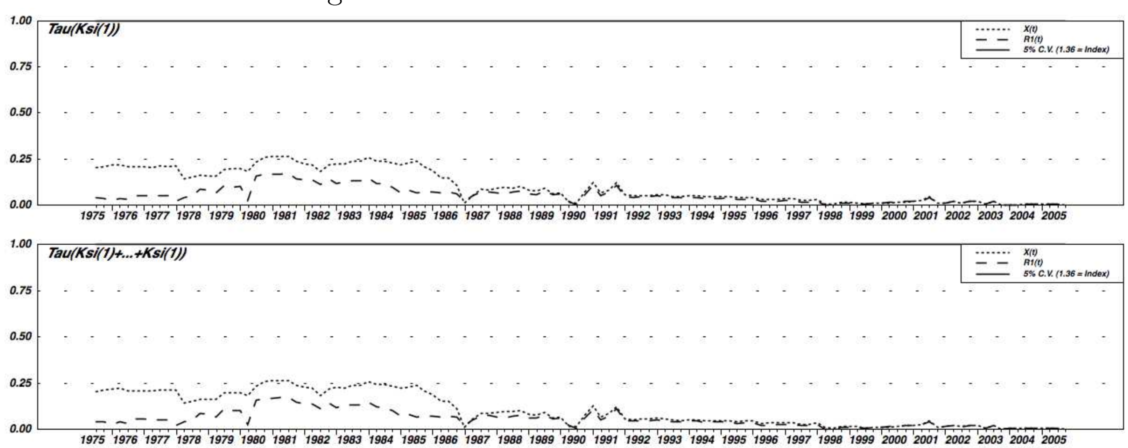

Figure 10: Max Test of a Constant Beta – MSI M1

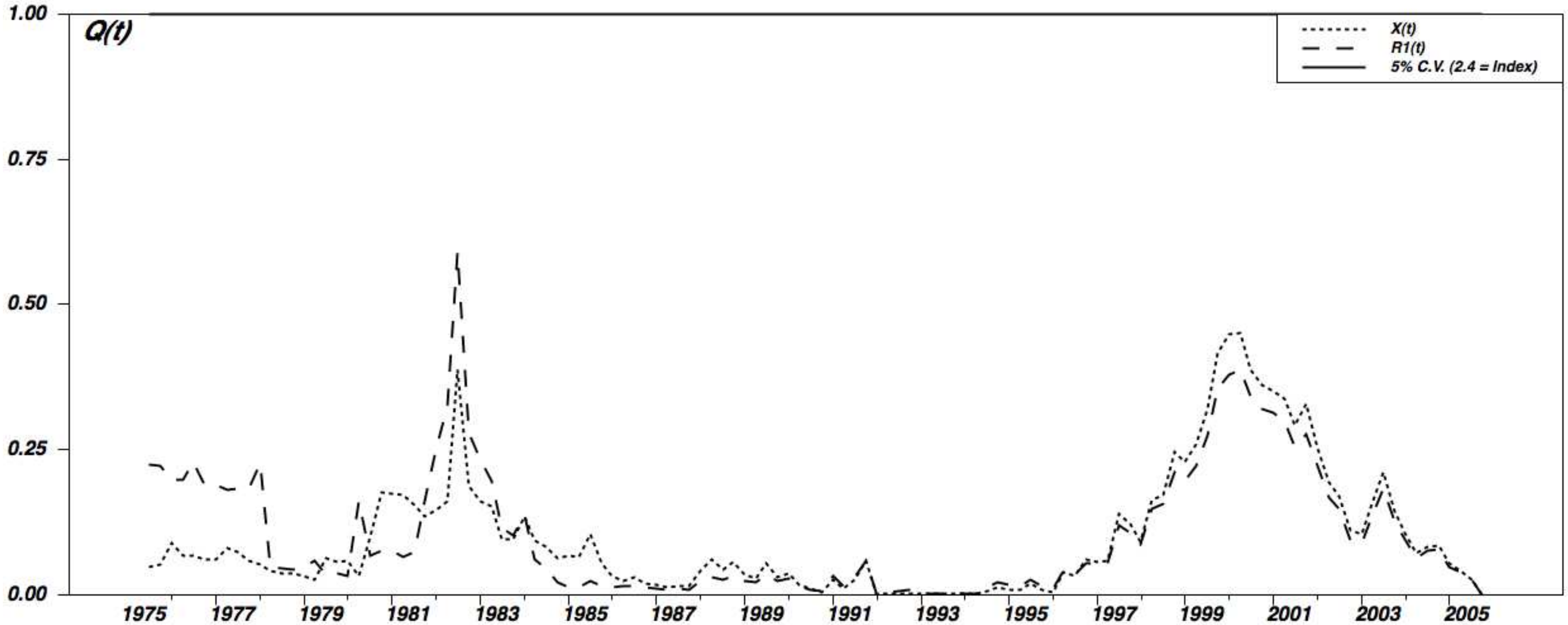

Figure 11: Max Test of a Constant Beta – MSI M2

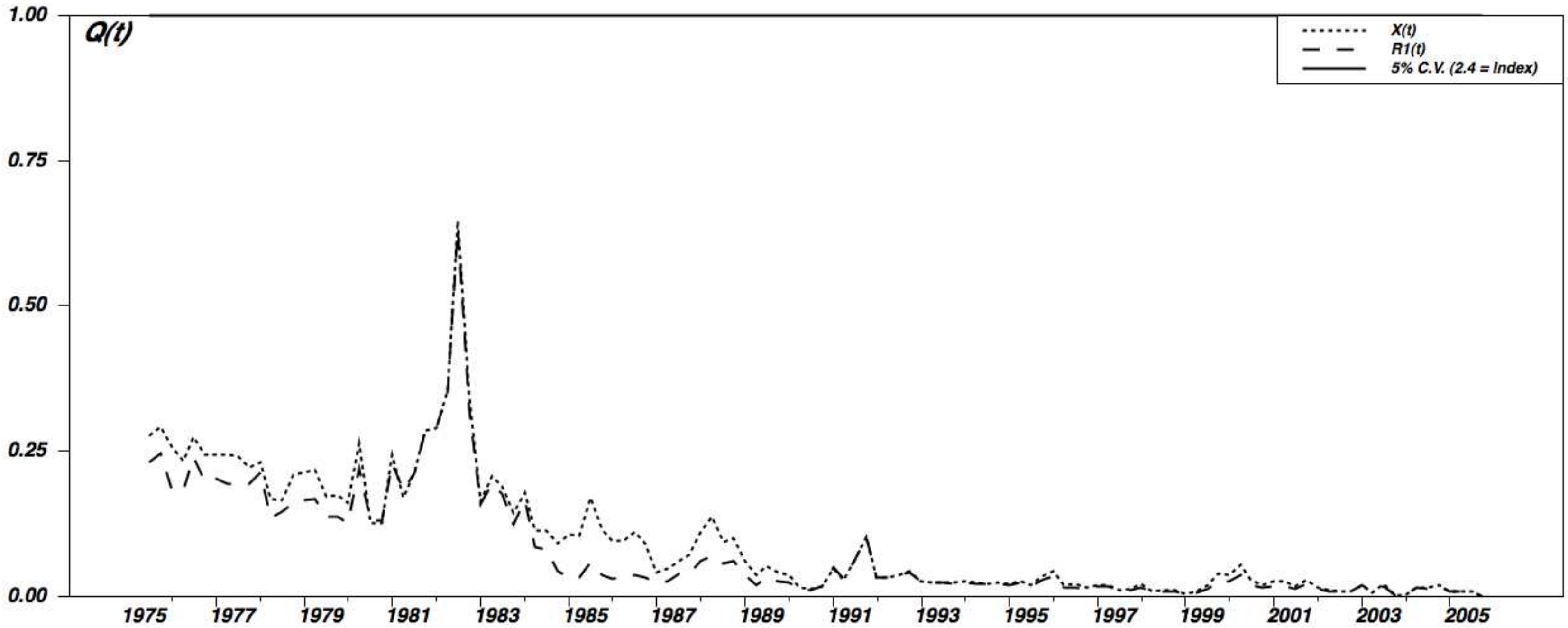

Figure 12: Max Test of a Constant Beta – MSI MZM

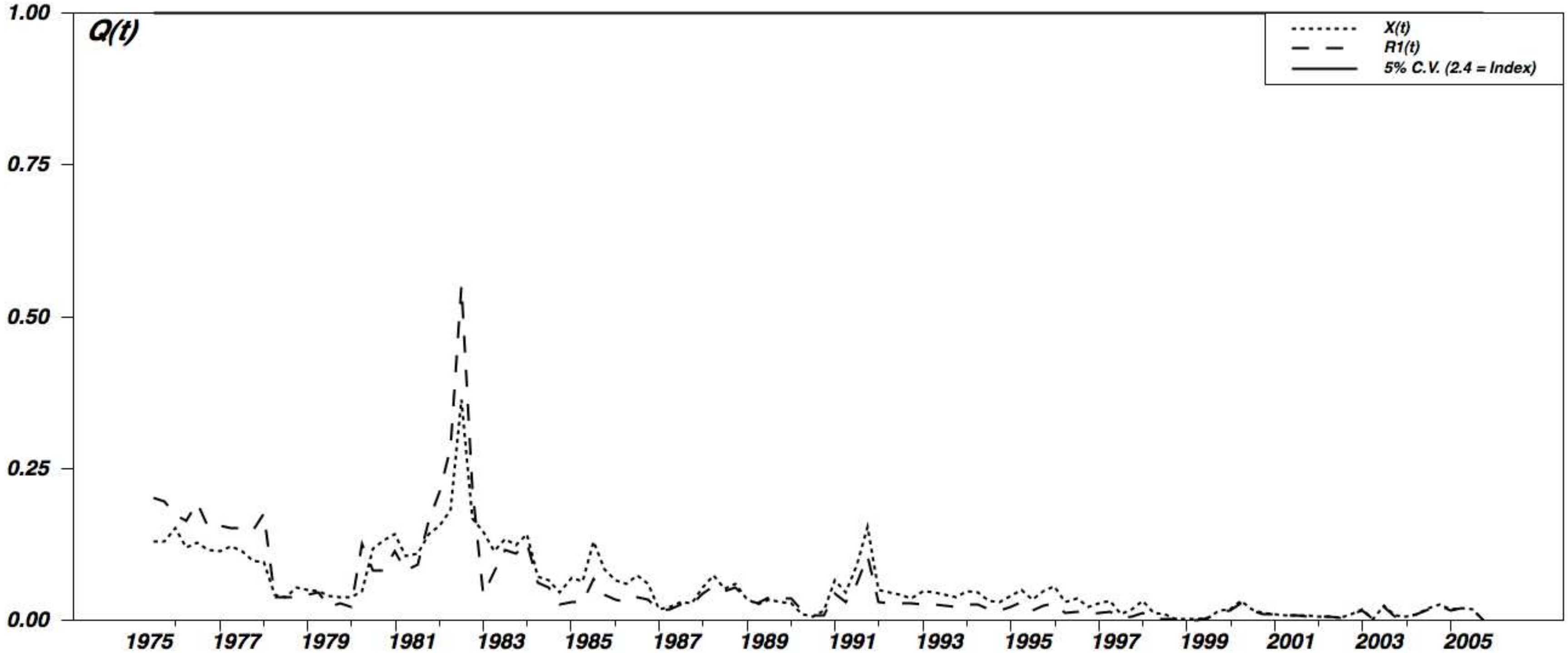

Table 3: Granger Causality Tests – MSI M1 Pre-1979

Variable	Nom. GDP	Inflation	MSI M1
Nom. GDP	0.81	0.82	0.23
Inflation	0.39	0.00	0.01
MSI M1	0.09	0.01	0.00

Table 4: Granger Causality Tests – MSI M2 Pre-1979

Variable	Nom. GDP	Inflation	MSI M2
Nom. GDP	0.64	0.73	0.09
Inflation	0.34	0.00	0.01
MSI M2	0.17	0.01	0.00

Table 5: Granger Causality Tests – MSI MZM Pre-1979

Variable	Nom. GDP	Inflation	MSI MZM
Nom. GDP	0.60	0.78	0.07
Inflation	0.36	0.00	0.02
MSI MZM	0.09	0.02	0.00

Table 6: Granger Causality Tests – MSI M1 Post-1979

Variable	Nom. GDP	Inflation	MSI M1
Nom. GDP	0.00	0.94	0.31
Inflation	0.15	0.00	0.02
MSI M1	0.61	0.02	0.00

Table 7: Granger Causality Tests – MSI M2 Post-1979

Variable	Nom. GDP	Inflation	MSI M2
Nom. GDP	0.00	0.45	0.31
Inflation	0.48	0.00	0.00
MSI M2	0.06	0.25	0.00

Table 8: Granger Causality Tests – MSI MZM Post-1979

Variable	Nom. GDP	Inflation	MSI MZM
Nom. GDP	0.00	0.25	0.19
Inflation	0.65	0.00	0.02
MSI MZM	0.00	0.31	0.00

Table 9: Central Bank Reaction Function – Post-1979

	Variable	Coefficient	t-stat
MSI M1	Constant	0.01	3.55
	ΔNGDP Forecast	0.03	0.62
MSI M2	Constant	0.02	7.28
	ΔNGDP Forecast	-0.06	-1.66
MSI MZM	Constant	0.03	7.20
	ΔNGDP Forecast	-0.20	-2.35

Table 10: Granger Causality p-values – MSI M1

Variable	HK Nom. GDP	HK Inflation	MSI M1
HK Nom. GDP	0.00	0.01	0.02
HK Inflation	0.03	0.03	0.04
MSI M1	0.17	0.39	0.00

Table 11: Granger Causality p-values – MSI M2

Variable	HK Nom. GDP	HK Inflation	MSI M2
HK Nom. GDP	0.00	0.07	0.02
HK Inflation	0.33	0.46	0.92
MSI M2	0.17	0.58	0.00

Table 12: Granger Causality p-values – MSI MZM

Variable	HK Nom. GDP	HK Inflation	MSI MZM
HK Nom. GDP	0.00	0.19	0.15
HK Inflation	0.02	0.67	0.68
MSI MZM	0.20	0.08	0.00

Table 13: IS equations

	RS	New Estimates							
		Cashless		MSI M1		MSI M2		MSI MZM	
Variable	1961 - 1996	1961 - 2005	1979:4 - 2005:4	61 - 05	79:4 - 05:4	61 - 05	79:4 - 05:4	61 - 05	79:4 - 05:4
Output gap (-1)	1.16	1.22***	1.25***	1.20***	1.25***	1.15***	1.24***	1.18***	1.25***
	(14.50)	(17.46)	(13.29)	(17.00)	(13.24)	(16.67)	(13.60)	(17.28)	(13.81)
Output gap (-2)	-0.26	-0.29***	-0.31***	-0.26***	-0.30***	-0.22***	-0.31***	-0.23***	-0.29***
	(-3.25)	(-4.12)	(-3.31)	(-3.61)	(-3.25)	(-3.19)	(-3.44)	(-3.41)	(-3.27)
Real Rate (-1)	-0.09	-0.07***	-0.05**	-0.07***	-0.04*	-0.05**	-0.03	-0.05**	-0.02
	(-2.75)	(-2.85)	(-2.06)	(-2.73)	(-1.83)	(-2.01)	(-1.44)	(-2.22)	(-0.96)
Money (-1)	–	–	–	0.10*	0.02	0.23***	0.19***	0.12***	0.10***
	–	–	–	(1.67)	(0.36)	(3.80)	(2.71)	(3.43)	(2.98)
$\bar{R}^2$	0.90	0.91	0.93	0.91	0.93	0.91	0.94	0.91	0.94
Durbin-Watson	2.08	2.15	2.12	2.16	2.12	2.19	2.15	2.19	2.15

***1% crit. value, **5%, *10%