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Summary

Changes in Monetary Policy Rules for the U.S.A., the U.K., and Japan: Estimating their Postwar Experiences

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This paper estimates monetary policy rules of postwar U.S.A., U.K., and Japan for three subperiods in each country. As was discussed by J.B. Taylor and M. Woodford, the rules can be broadly called the Taylor rule. Using the GMM and related method, we obtain satisfactory estimation results. We also compare the size of policy rate responses to three explanatory variables, which are expected inflation, the output gap, and foreign exchange reserves among the three countries.

Article

Changes in Monetary Policy Rules for the U.S.A., the U.K., and Japan: Estimating their Postwar Experiences

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1. Introduction

This paper estimates monetary policy rules of the three countries for postwar three periods, covering around the 1960s through early 2010s. Dividing the whole half a century into three periods for each country will enable us to evaluate changes in policy rules for each country, and also compare the policy rules in cross-country manners. We use the generalized method of moments (GMM) estimation and related method for this purpose. The divisions into three periods will be made referring to historical episodes as well as to breakpoint tests of Chow and Quant-Andrews, and the Cusum-of-squares tests.

As for policy rules, we consider the well-known Taylor rule (Taylor 1993) as a baseline formulation. As is pointed out so far, one does not have clear evidence that the countries we are concerned with here have followed the Taylor rule. Taylor argues that if the country exhibits stable, successful macroeconomic performance, their monetary policy can be approximated by his rule; see Woodford (2001, 2003) and Taylor et al. (2010). The difficulty for policymakers to follow his rule and for analysts to assimilate policy practice with the rule in a more or less exact manner will be that his rule involves two unknown time-dependent parameters: the natural (equilibrium) real interest rate and the gap in aggregate demand and supply for output, both of which are of ex-ante nature.

In this context, Trehan and Wu (2007) present a model based on Kalman-filtering, and estimates natural (equilibrium) interest rates. Curdia et al. (2015) interpret the policy process of the Greenspan period to be of natural interest rate-targeting, assuming that the criteria for manipulating policy rates are based on the inflation target as well as the gap between the observable interest rate and efficient rate, the latter of which is a concept similar to Wicksell's natural rate.

Judd and Rudebusch (1998) presuppose the Taylor-type rate-setting equation (the policy reaction function), and divide their whole period of analysis into three periods, which is based on the periods of Fed Reserve Board chairmanship of Burns, Volker, and Greenspan. Then they compare the three policymaking regimes in terms of differences in coefficients of reaction functions.

Clarida, Gali, and Gertler (1998, 2000) estimate Taylor-type policy reaction functions for several countries in 1998 and for the U.S. in 2000, and in the former paper, they suggest the main concern of the countries they dealt with is stabilization of inflation. In the latter paper, they show that after Burns's chairmanship of the Fed, the policy turned more aggressive in fighting inflation and brought about higher economic stability

Estrella and Fuhrer (2003) suppose policy reaction functions, which are expressed in differenced form in the policy rate and are similar in concept to the Taylor rule. Then, based on breakpoints which are detected using some breakpoint tests, they examine the stability of the reaction functions before and after the breakpoints, and compare the stability of reaction functions, among others, when inflation expectations are forward-looking or backward-looking.

Zhang, Osborn, and Kim (2008), focusing on the U.S. Phillips curve for the post-1970 period, examine if the whole period had structural changes (breakpoints) once or twice. This study is done using some tests for structural change, and shows, among other points, and similarly to Estrella and Fuhrer (2003), that the methods they employ are not necessarily unanimous in deciding the breakpoints (dates of structural change).

In this paper we deal with the three countries, the U.S.A., the U.K., and Japan, and suppose that the Taylor rule approximates the monetary policymaking in those countries.

A novel point coming out of the following analysis is that, for those three countries, in eight out of nine periods, the levels of foreign exchange reserves affected their policy-rate settings. We will also show that the policy rate responses to the variables the central banks are concerned with exhibit fairly noticeable differences among the three countries we deal with.

The plan of the paper is as follows. Section 2 describes the model, and baseline estimations using the OLS to ascertain two borderline periods for each country. Then for three periods of each country, we estimate the Taylor rules which assimilate the actual policymaking processes, using the GMM and related method with own explanatory variables as instruments. Section 3 concludes.

2. Piecewise Policy Estimation: The Baseline Cases

We start with a brief description of the Taylor rule, mainly based on Clarida, Gali, and Gertler (2000). The rule can be written as

$$r_t^* = r^* + \beta(E\pi_{t,k} - \pi^*) + \gamma E(x_{t,q}) + \varepsilon_t, \tag{1}$$

where r_t^*: the nominal target policy rate (the federal funds rate for the U.S.A., the treasury bill rate for the U.K., and the call rate for Japan); r^*: the natural (equilibrium) nominal interest rate equal to Wicksell's natural rate; π^*: the target inflation rate; $E\pi_{t,k}$: the inflation rate expected in period t for period k ($k > t$); $E(x_{t,q})$: excess demand for output expected in period t for period q; β, γ: positive constants; E: the expectation operator; and ε_t: a disturbance term. To simplify analysis we assume that policymakers use x_t instead of $E(x_{t,q})$ in their rate setting (in other words, they set the rate approximating x_t to $E(x_{t,q})$).

Also, Woodford (2003), among others, shows that, along with an inflation equation (Phillips curve equation), the policy rule (1) can be derived from the cost minimization of policymakers, the cost being the weighted sum of squared deviations of inflation and the output gap from the targets of two variables.

For the U.S. economy, we divide the whole (effective) estimation period into 1960：1 (the first quarter of 1960) through 1979：4, 1980：1 through 2001：4, and 2002：1 through 2013：2. It was in 1979：4 that Paul Volker took office of the Fed Reserve Board chairman. The period 2001：4 was around the bottom of the business cycle, and growth rates of real GDP over past four periods (decimal values) are 2001：1＝0. 019, '01：2＝0. 006, '01：3＝0. 003, '01：4＝0. 002, and '02：1＝0. 010. Data were taken from the International Financial Statistics of the International Monetary Fund.

To ascertain if the above period divisions are supported by statistical tests, we resort to Chow's breakpoint test (Chow 1960), Quant-Andrews' breakpoint test (Andrews 1993), and the Cusum-of-squares test (Johnston and DiNardo 1997); see also Maddala and Kim (1998, Ch. 13), who suggest using casual, historical observations in conjunction with statistical tests in deciding the dates of structural change. To apply Chow and Quant-Andrews tests, it is necessary to get an OLS estimation of the rule for the whole period, 1960 : 1 through 2013 : 2. In deriving the output gap, we use either the Hodrick-Prescott (H-P) business cycle factor (logged output minus the trend factor in the H-P filter) or estimated series of output on time trend and trend squared. Since in this estimation of the whole period, the second output gap did not take on a significant positive coefficient, I only used an H-P filtered gap.

We denote *ffa*: the federal funds rate (a dependent variable; the last letter *a* implies it is the U.S.'s), *egpa*: expected inflation of the GDP deflator, *qgpha*: H-P filtered output gap. *egpa* is given by the estimated value of inflation rate of the next period, which is regressed on the current and one-period past inflation rates.[1)] Also, we include the possibility of policymakers to consider the level of foreign exchange reserves (*lira*) in rate setting. Though *lira* does not appear as a significant variable in the whole period estimation, most of the piecewise estimations for the three periods of the three countries show the foreign exchange reserves to be significant variables. The result turns out, for period 1960 : 1 through 2013 : 2, as:

	ffa	*egpa*	*qgpha*	*lira*
coefficient	4.836	0.320	−0.001	
p-value	0.000	0.000	0.338;	$r^2 = 0.421$, $DW = 0.257$

The intercept of the regression was present in estimation but is omitted above. r^2 is the coefficient of determination adjusted for the degree of freedom. The low DW-ratio reflects a small number of explanatory variables for a long range of period. Data for foreign exchange reserves (logged, in million dollars; the units of the U.K. and Japanese counterparts are the same) were taken from the Main Economic Indicators of the OECD and the International Financial Statistics (IFS) of the International Monetary Fund;[2)] all the other data used in this paper draw on the IFS.

In Chow's test, if one posits the break dates to be 1980 : 1 and 2002 : 1, p-values of the F-statistic, log likelihood ratio and Wald statistics are all zero under the null of no breaks, which support our break points posited above.

Quant-Andrews' test, which concerns unknown breakpoints, detects the date 2001 : 4 only, and for this date, three pairs of LR F-statistics and Wald F-statistics, each including maximum, exponential, and average statistics, all show p-values of zero under the null of no breakpoints. The test, however, failed to detect our another breakpoint '80 : 1, although the dip of real GDP growth rates was deeper around '80 : 1 than around '01 : 4.

Though I omit the graphical presentation of the Cusum-of-squares (COS) test, the COS line stays out of the stability line (5% significance level) at our breakpoints, which also supports our supposed dates (see Johnston and DiNardo 1997 for the test).

We first turn to the U.S.'s first period ('60 : 1–'79 : 4). Anticipating the total nine cases for the three countries, the output gaps using the H-P filtered gap yield better results in all the (nine) cases, so that I will exhibit the results with the latter measure. Most of the estimations that follow use the generalized method of moments (GMM, HAC) with lagged values of explanatory variables as instruments. HAC is the estimation weighting matrix for this case.[3] The last letter *a* attached to the three variables implies country attribution; letter *b* implies the variables belong to Great Britain, and *j* to Japan.[4]

ffa	*egpa*	*qgpha*	*lira*	
coefficient	0. 817	0. 602	−0. 009	
p-value	0. 000	0. 000	0. 022;	$r^2 = 0.631$, *prob* (*J-stat*) = 0. 151 (2)
inst. lag	2	2	2	

where *qgpha* is the business cycle factor of the Hodrick-Prescott filter (the difference between logged output and the trend factor). The row 'inst. lag' exhibits the number of instruments used. For example, 2 below *egpa* means that *egpa*(-1) and *egpa*(-2) served as the instruments. In the above, the overidentified restriction is satisfied because it is larger than 0. 05, which means that the instruments are likely to be properly chosen. See, e.g., Hamilton (1994, Ch. 14) or Davidson and MacKinnon (2004, Ch. 9) for the overidentifying restrictions and J-statistics.

Using the GMM (White), where White means the weighting matrix for this case, the U.S. second period ('80 : 1–'01 : 4) yields

ffa	*egpa*	*qgpha*	*lira*	
coefficient	1. 030	0. 522	0. 023	
p-value	0. 000	0. 001	0. 000;	$r^2 = 0.784$, *prob* (*J-stat*) = 0. 784 (3)
inst.lag	2	2	1	

Compared to U.S.'s other periods, the last of which will be shown shortly, the coefficients on expected inflation exceeds unity in this period, so the interest rate policy has been stable and countercyclical in this period; see Clarida, Gali, and Gertler (1998, 2000), and Woodford (2001).

In the third period ('02 : 1–'13 : 2) of this country, the very low interest rate policy has been adopted since the end of 2008, but with the GMM (White), the estimation turns out satisfactory:

ffa	*egpa*	*qgpha*	*lira*		
coefficient	0.954	0.462	−0.031		
p-value	0.007	0.028	0.000;	$r^2 = 0.757$, *prob* (*J-stat*) $= 0.275$	(4)
inst. lag	2	2	1		

Note that in view of the coefficient on expected inflation, the rate setting was not done in a stable manner.

Turning next to the estimation of the U.K., we first derive the policy rule relationship for the whole period, 1960 : 1 through 2014 : 1, using the OLS, which reads as

trb	*egpb*	*qgphb*	*lirb*	
coefficient	3.740	0.511	0.000	
p-value	0.000	0.000	0.843;	$r^2 = 0.318$, $DW = 0.198$

where *trb* is the treasury bill rate which the Bank of England is concerned with.

We will posit the three sub-periods as '60 : 1–'75 : 1, '75 : 2–'00 : 4, and '01 : 1–'14 : 1. Chow's breakpoint test supports this divisions, and the COS test shows that the two breakpoints ('75 : 1 and '01 : 1) stay in the unstable region of regression coefficients. But Quant-Andrews' test indicates one breakpoint, which is '06 : 1. However, assuming this breakpoint, probably because of few sample points in the third period, the coefficient of *egpb* turns out insignificant ($p = 0.585$), though '06 : 1 belongs to an unstable coefficient region in the COS test. For those reasons, we posit the second breakpoint to be '01 : 1.

The U.K.'s first period ('60 : 1–'75 : 1) was estimated using the limited information maximum likelihood method as

trb	*egpb*	*qgphb*	*lirb*		
coefficient	0.707	2.197	−0.006		
p-value	0.008	0.002	0.450;	$r^2 = -1.275$	(5)
inst.lag	1	2	2		

The value of r^2 is not important in the instrumental variable method; see, e.g., Wooldridge (2014, Ch. 15).

The second period ('75 : 2–'00 : 4) of this country yields, using the GMM (White), the following result:

trb	*egpb*	*qgphb*	*lirb*		
coefficient	0.270	0.639	−0.010		
p-value	0.002	0.021	0.038;	$r^2 = 0.276$, *prob* (*J-stat*) = 0.135	(6)
inst.lag	2	1	2		

The third period ('01 : 1–'14 : 1) of the U.K. produces the result, using the GMM (HAC), that

trb	*egpb*	*qgphb*	*lirb*		
coefficient	0.099	0.102	−0.046		
p-value	0.062	0.034	0.000;	$r^2 = 0.802$, *prob* (*J-stat*) = 0.641	(7)
inst. lag	2	2	2		

We finally deal with the Japanese case. The estimation of the reaction function for the whole period (1959 : 3–2013 : 1) uses the H-P filtered output gap *qgphj* for the thrust on output demand.

crj	*egpj*	*qgphj*	*lirj*	
coefficient	0.339	0.112	−0.010	
p-value	0.000	0.097	0.000;	$r^2=0.806$, $DW=0.242$

where *crj* stands for the call rate (short term interbank lending rate). For Japan we posit two breakpoints, 1974：1 and 1991：1; the former is the next quarter of the first oil-supply shock, which is generally regarded as a major factor for Japan to terminate the 'high-growth period,' while the second breakpoint represents the date when the 'bubble period' and 'middle-size growth period' ended, and it entered the 'low-growth period.' For those tentative breakpoints, Chow's test supports them in terms of *F*-statistics, log-likelihood, and Wald-statistics (their *p*-values are zero).

Referring to Quant-Andrews' unknown breakpoint test, it indicates 1973：3 and 1984：1; the former date is one-period preceding the first oil shock (it occurred in '73：4), while the latter is one-year earlier than the start of the bubble period. Though this test suggests different dates, using '84：1 makes our three periods with quite unbalanced lengths, so that we will adopt the date divisions we proposed initially.

Denoting the county attribution by *j*, Japan's first period ('59：3–'73：4) yields, using the GMM (HAC):

crj	*egpj*	*qgpj*	*lirj*		
coefficient	0.190	0.262	−0.006		
p-value	0.003	0.055	0.045;	$r^2=261$, $prob\,(J\text{-}stat)=0.576$	(8)
inst. lag	2	2	2		

Here as the output gap, we use the residual in the regression of logged output on time-trend *tr* and its squared series with $tr\,(1)=1$ at 1975：1. When the output gap is represented by H-P filtered *qgphj*, its

p-value is 0. 477.

The reaction function for the second period ('75 : 1–'90 : 4) was estimated with the GMM (HAC) as

crj	*egpj*	*qgphj*	*lirj*	
coefficient	0. 423	1. 203	−0. 015	
p-value	0. 037	0. 006	0. 000;	$r^2=0.390$, $prob\,(J\text{-}stat)=0.662$ (9)
inst.lag	2	1	2	

Finally, Japan's third period ('91 : 1–'13 : 1) turned out with the GMM (White) as

crj	*egpj*	*qgpj*	*lirj*	
coefficient	0. 336	0. 076	−0. 009	
p-values	0. 001	0. 006	0. 000;	$r^2=0.579$, $prob\,(J\text{-}stat)=0.550$ (10)
inst.lag	1	1	2	

As in the previous period, the excess demand for products is represented by detrended output. If instead we use the business cycle factor of the Hodrick-Prescott filter *qgphj*, its coefficient has $p=0.864$.

It is in order here to compare among the three countries the sizes of coefficients for each of the three explanatory variables. If we write the average of coefficients of the U.S.'s expected inflation rates $egpa^*$, with the U.K. and Japanese counterparts $egpb^*$ and $egpj^*$, respectively, we have

$egpa^*=0.934$, $egpb^*=0.358$, $egpj^*=0.246$.

Similarly, the average coefficients of output gaps for the three countries, $qgpa^*$, $qgpb^*$, and $qgpj^*$, become

$qgpa^*=0.529$, $qgpb^*=0.979$, $qgpj^*=0.538$,

Also, the averages of the effects on rate settings of foreign exchange reserves are

$lira^* = -0.021$, $lirb^* = -0.028$, $lirj^* = -0.010$,

where the first-period *lirb* is excluded because the coefficient is not significant.

From the above statistics one can get the following observations:

(i) The effect of expected inflation on the rate settings is largest in the U.S., the U.K. comes the second, and Japan's is the smallest. The U.S.'s average is about three times those of the other two countries.

(ii) The output gaps affect rate settings to the largest degree in the U.K., while the averages of the U.S. and Japan are of the similar size, which are about half of the U.K.'s.

(iii) The effects of foreign exchange reserves (in absolute value) are the largest in the U.K., and then come the U.S. and Japan. The average of the U.K. is about three times that of Japan.

3. Conclusions

This paper has examined the monetary policy reaction functions of the three countries, the U.S.A., the U.K., and Japan, by dividing the whole period of analysis into three sub-periods. Policy reaction functions were supposed to be based on the Taylor rule. Although Taylor himself has shown that his rule describes the policy process of industrial countries reasonably well, since this rule involves two unknown variables, the (real or nominal) natural interest rate and the output gap, it will be a necessary and worthwhile step to estimate policy reaction functions in time-series (historical) as well as cross-country manners.

As for the output gap, we suppose two alternatives: the Hodrick-Prescott filter and the residuals in the OLS of logged output on the quadratic time trend.

The main findings from our inquiry would be first that, except for the U.K.'s first period, the level of foreign exchange reserves had negative impacts on the policy rate setting.

The second point to be noted is that the policy rate responded to higher expected inflation in the U.S.A. to the largest degree, which is about three times that of the U.K. and Japan.

The third point having emerged from our study is that the response of policy rates to the level of foreign exchange reserves, in absolute value, was the largest in the U.K., the second being the U.S.A., and Japan was the smallest, where the size of the U.K. is about three times that of Japan.

As is well known recently Taylor et al. (2010), because of the U.S. financial market turmoil starting a few years prior to 2010, the policy rates of the U.S. and the U.K. were stuck near the zero floor. Also, in Japan the matters turned into a similar situation since around the turn of the century. Hence the description of monetary policymaking during these developments may need extra consideration, which will, however, be left for another occasion.

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Notes

1) The next-period inflation rate *gpa*(1) is regressed on a constant, *gpa* and *gpa*(-1). The estimated inflation rate is expected inflation *egpa* (for *gpa* of the next period). Since the residual of the above estimation can be regarded as *I*(0), i.e. covariance-stationary, *egpa* is a rational expectation.

2) Since $\partial ffa/\partial lira$ is a semi-elasticity, where *ffa* is in decimal numbers, measuring units of foreign exchange reserves do not affect the estimation.

3) From the two weighting matrixes HAC and White, I adopted the one with better *p*-values and *Prob*(*J-stat*) in each case.

4) Concerning the policy rule estimation for the three countries with output gaps measured by quadratic trends or Hodrick-Prescott filters, I did not include as explanatory variables, policy rates with one-period or two period lags, which represent policy smoothing or policy inertia. The reasons for this include that smoothing will occur in shorter time span like weekly or monthly bases, that various shocks did occur on rate-setting, and that data are recently revised one rather than real-time one, as was described by Rudebusch (2002).

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