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Exchange Rate Regimes and Inflation Persistence

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Some empirical research has suggested that inflation is more persistent under floating exchange rates. Theoretically, we should expect a higher variance of inflation persistence across countries under floating rates, but not necessarily a higher mean. It is shown that estimates of inflation persistence are biased upwards by underfitting mean shifts in the sample. After correction for mean shifts, there is evidence of greater inflation persistence in the post–Bretton Woods period, but no evidence of variation across exchange rate regimes. Monetary growth has been much less accommodative of inflation since 1979, with no difference between EMS and non-EMS countries. [JEL E31, E50, F31, F41]

Does the exchange rate regime affect how the monetary authorities react to inflation shocks and, consequently, the persistence of those shocks? Intuitively, one would answer in the affirmative, in view of the powerful constraints on domestic policy actions imposed by a commitment to a fixed exchange rate. Several empirical studies have suggested that inflation has been significantly more persistent (positively serially correlated) and monetary policy more accommodative of inflation under floating exchange rates than under pegged rates (see Alogoskoufis and Smith, 1991; Alogoskoufis, 1992; Obstfeld, 1995). These results have been questioned recently by Burdekin and Siklos (1999), who argue that, though there have been historical shifts in inflation persistence, they have not been closely associated with exchange rate regime changes. The general issue was first raised by Dornbusch (1982) in the context of a country operating a linear policy rule for exchange rate adjustment

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in a pegged-rate regime. Dornbusch showed, the more that the exchange rate accommodated inflation shocks, the more persistent these shocks were.

None of these papers fully explores the theoretical dimensions of the problem. Dornbusch (1982) treats the policy rule as exogenous, and does not consider floating exchange rate regimes. Adams and Gros (1986) and Montiel and Ostry (1991) examine the implications of targeting the real exchange rate. Alogoskoufis and Smith (1991), Alogoskoufis (1992), Obstfeld (1995), and Burdekin and Siklos (1999) provide a purely empirical analysis of the impact of exchange rate regimes.

The first aim of this paper is therefore theoretical. I analyze the dynamics of inflation as chosen by a government maximizing a social welfare function, which includes a term in the deviation of domestic from foreign inflation. The weight attached to this element of the loss function varies with the exchange rate regime, as a way of capturing the loss of monetary independence associated with pegging the exchange rate.

On the empirical level, the evidence for OECD countries is apparently rather striking. Simple regressions suggest that, under floating rates, inflation is more persistent and monetary policy is much more likely to accommodate an inflation shock. Alogoskoufis and Smith (1991) estimate first-order autoregressive coefficients for the annual change in the GDP deflator for the United States, and their point estimate is 0.27 for the period 1948–67 and 0.70 for the period 1968–87. Their results for the United Kingdom are very similar. Alogoskoufis (1992) extended this work on inflation persistence to 21 OECD countries. For the OECD weighted average, his results are very similar to those of Alogoskoufis and Smith. For the individual countries, he finds that persistence increased in 20 out of 21 cases, comparing 1972–87 with 1953–71. Obstfeld (1995) revisited the issue, using data for 12 OECD countries over the period 1953–94. Dividing the data into 1953–72 for fixed rates and 1973–94 for floating rates, he finds that persistence increased in the later period (although not always by much) for all countries except the United States, which had particularly high persistence in the fixed-rate period. Obstfeld attributes his findings to the reserve-currency role played by the United States under the Bretton Woods system. Finally, Burdekin and Siklos (1999) analyze long runs of annual data for Canada, Sweden, the United Kingdom and the United States, using a procedure for identifying and dating multiple structural breaks in a persistence regression. Their results suggest a statistically significant (but relatively small) upward shift in persistence some time between 1974 and 1981 in each country, although they question whether there is any historical association between shifts in persistence and exchange rate regime changes.[1]

In the empirical section of this paper, it is first shown that estimates of inflation persistence are highly sensitive to shifts in mean inflation during exchange rate regimes. The crucial point is that, if n mean shifts occur during an exchange rate regime, then allowing for $n - 1$ or fewer mean shifts biases upwards the

[1]The estimated shifts in persistence and dates are: Canada +0.03 (1981), Sweden +0.07 (1979), UK +0.11 (1974) and US +0.06 (1979).

estimate of persistence, whereas allowing for $n + 1$ or more shifts induces a downward bias of much smaller proportions.[2] This effect explains why dating the exchange rate regime change at 1968 (as in Alogoskoufis and Smith, 1991) yields much bigger differences in persistence than dating it later (as in Obstfeld, 1995): mean inflation shifted upwards in the later 1960s, and underwent several further shifts in later years, but was rather stable from 1950 to 1967. I reanalyze the postwar evidence on exchange rate regimes and inflation persistence for a number of OECD countries, allowing for this mean shift effect.

What about monetary policy under fixed and floating exchange rates? Using data back to 1872, Alogoskoufis and Smith (1991) report that weighted average monetary growth in the United States and the United Kingdom accommodated about half of inflation under floating rates, but did not accommodate it at all under fixed rates. Interpreted in terms of a monetary policy reaction function, this suggests that monetary policy was tightened much more in response to an inflationary shock under fixed than under floating rates. As far as the postwar period is concerned, this is a surprising finding, in view of the evidence from recent research that, since 1979, the Federal Reserve's real interest rate response to U.S. inflation shocks has been much stronger than before, presumably in recognition of expectational shifts in the Phillips curve (Clarida and others, 1998a; Taylor, 1999). My investigation of the dynamics of monetary growth supports the conclusion that there has been a much stronger monetary policy response to inflation shocks in recent decades. I find that monetary growth in the United States was strongly accommodative of immediate past inflation in the Bretton Woods period, but has been much less so under floating rates. Comparing floating-rate countries with members of the European Monetary System (EMS) for the 1980s and 1990s, I find no evidence of greater monetary accommodation under floating exchange rates.

I. Theory

Alogoskoufis and Smith (1991), Alogoskoufis (1992), and Obstfeld (1995) all work with variants of the same model, which assumes overlapping contracts, with each contract having a fixed probability of being renegotiated, along the lines of Taylor (1980).[3] If each contract has a probability $1–\psi$ of renegotiation in each period, then

$$p_t = \psi p_{t-1} + (1 - \psi) x_t \qquad 0 \le \psi \le 1 \qquad (1)$$

[2]As is discussed below, it matters whether or not we are concerned about bias relative to the true value of the parameter or merely relative to the estimate from another sample. It is the latter which is important in the present context, where comparisons of estimates are being made across exchange rate regimes. The point made here is not the same point as that made by Burdekin and Siklos, who test for *simultaneous* shifts in mean inflation *and* the persistence coefficient, so their analysis does not allow for mean shifts within any given period when persistence is assumed to be constant.

[3]Dornbusch (1982) uses a slightly different version of the model in which each contract lasts for precisely two periods.

where x represents the level of newly contracted prices and p is the price level (all in logs). Taking first differences of equation (1) shows that, unless new-contract inflation (Δx) is (sufficiently) negatively autocorrelated, aggregate inflation will be positively autocorrelated, since

$$\Delta p_t = (1-\psi)\Delta x_t + \psi(1-\psi)\Delta x_{t-1} + \psi^2(1-\psi)\Delta x_{t-2} + \ldots \tag{2}$$

Negative autocorrelation of Δx could, of course, easily arise if monetary policy refused to accommodate new-contract inflation. The next step is that the money supply (m) in each country (in logs) is assumed to be related to the current price level through a simple monetary policy reaction function:

$$\Delta m = \alpha + \beta \Delta p \tag{3}$$

In this equation, α is interpreted as a trend which reflects output growth and is independent of current inflation, and β represents the degree to which the authorities are willing to accommodate the current level of prices. In this model, the degree of persistence of inflation (and obviously, therefore, of monetary growth) increases with the accommodation parameter, β (see the appendix to Obstfeld, 1995 for details). Alogoskoufis and Smith (1991), Alogoskoufis (1992), and Obstfeld (1995) all speculate that β is likely to be higher under floating rates than under fixed rates.[4] In the empirical implementation, because of lack of data on x, the procedure adopted by these authors is to estimate a first-order autoregression for inflation, allowing the coefficients to change across exchange rate regimes, and to estimate equation (3), again allowing for coefficients to change across regimes.

In order to address the theoretical issue of inflation persistence under different exchange rate regimes, I employ a model developed from the stochastic version of the Barro-Gordon (1983) macroeconomic policy game introduced by Rogoff (1985). In the closed-economy version of the model, the government maximizes a utility function defined over inflation (π) and output (y), where Y represents equilibrium output:

$$Z = -0.5\pi^2 - 0.5b(y - Y - k)^2 \qquad b,k > 0 \tag{4}$$

In this equation, b represents the relative weight attached to the output objective, while k is a device for introducing inflationary bias into the model. To this we add an expectations-augmented Phillips curve that is subject to stochastic shocks:

$$y = Y + a(\pi - \pi^e) + \varepsilon \qquad a > 0 \tag{5}$$

where ε is a supply-side shock which is assumed to follow the process:

[4]The use of the word "speculate" is deliberate. These authors do not derive this prediction from their theoretical model; indeed Dornbusch makes this very point in his comments on Obstfeld's paper (Obstfeld, 1995, p. 200).

$$\varepsilon_t = c\varepsilon_{t-1} + \eta_t \qquad \eta \sim N(0,(1-c^2)\sigma^2 I) \tag{6}$$

The parameter c (assumed > 0) determines the degree of autocorrelation in supply-side shocks. The variance of η varies with c so as to standardize the variance of ε at σ^2. In the context of the model described by equations (4) and (5), equation (6) is a simpler way of introducing inflation persistence than through the overlapping contracts model of equation (1).

The government chooses the inflation rate in order to maximize equation (4), after observing the current-period shock. The private sector, however, is assumed to form its expectations before observing the current-period shock. Substituting for y from equation (5), differentiating equation (4) and setting the differential equal to zero yields the following:

$$(1 + a^2b)\pi = a^2b\pi^e + ab(k - \varepsilon) \tag{7}$$

Assuming rational expectations, the private sector's inflationary expectations are determined by taking the mathematical expectation of equation (7). Using the fact that, from equation (6), $E(\varepsilon_t) = c\varepsilon_{t-1}$, this yields

$$\pi_t^e = ab(k - c\varepsilon_{t-1}) \tag{8}$$

In other words, inflationary expectations embody information about the expected shock in time t conveyed by the observed shock in time $t - 1$. Substituting from equation (8) back into equation (7) yields

$$\pi_t = ab(k - c\varepsilon_{t-1}) - \left[(ab)/(1 + a^2b)\right]\eta_t. \tag{9}$$

Within the context of this model, we may define monetary accommodation as the degree to which a supply-side shock is absorbed in price rather than in output fluctuations. From equation (9), it is evident that the anticipated component of the shock ($c\varepsilon_{t-1}$) is more strongly accommodated than the unanticipated component (η_t). Since c determines the proportion of the shock which is anticipated, it follows that more persistent shocks are more strongly accommodated. Equation (9) shows that the unconditional mean of inflation is abk and, since we may write, using the fact that from equation (6) $c\varepsilon_{t-2} = \varepsilon_{t-1} - \eta_{t-1}$,

$$\pi_{t-1} = ab\left(k - \varepsilon_{t-1}\right) + ab\eta_{t-1} - \left[(ab)/\left(1 + a^2b\right)\right]\eta_{t-1} \tag{10}$$

it follows that the covariance of π_t and π_{t-1} is given by

$$\mathrm{cov}\left(\pi_t, \pi_{t-1}\right) = c(ab)^2\sigma^2 \tag{11}$$

The variance of π_t is:

$$\mathrm{var}\left(\pi_t\right) = (ab)^2\sigma^2\left[c^2 + \left(1 + a^2b\right)^{-2}\left(1 - c^2\right)\right] \tag{12}$$

from which it follows that the correlation between π_t and π_{t-1} (i.e., inflation persistence) is:

$$\mathrm{corr}(\pi_t, \pi_{t-1}) = c\Big/\left[c^2 + (1 + a^2 b)^{-2}(1 - c^2)\right]. \tag{13}$$

This is zero when $c = 0$, and increases with c to reach a maximum at $c^2 = 1/[(1 + a^2b)^2 - 1]$, which could be greater or less than one. Inflation persistence either increases monotonically with c, or increases with c up to this maximum, and then declines.

So far, this is a model of a closed economy. The only difference from Rogoff (1985) lies in the assumption of persistence in the supply-side shock. To derive an open-economy version capable of allowing for different exchange rate regimes, I introduce an additional term in the government's utility function that reflects the costs of deviating from the inflation rate of the foreign country which issues the reserve currency of the fixed-rate system. These costs will reflect the exchange rate regime in operation. Thus equation (4) becomes

$$Z = -0.5\pi^2 - 0.5b(y - Y - k)^2 - 0.5h(\pi - \pi^*)^2 \qquad h \geq 0 \tag{14}$$

where π^* is the inflation rate of the reserve currency. The idea is that, under a system of pegged exchange rates, h will be large and the last term will dominate the other terms in equation (14), inducing the government to choose an inflation rate close to π^*. The costs of deviating from π^* will be reflected in a balance of payments disequilibrium and the policy adjustments required to correct it. Because of nontraded goods, transport costs, etc., it is realistic to permit some deviation of π from π^* (i.e., f is not infinite). Under floating rates, there is no cost to deviating from $\pi = \pi^*$, since the exchange rate is free to adjust, so that $h = 0$.

Pegged exchange rates are assumed to impose no constraints on the reserve country, which consequently chooses its inflation rate as if it were a closed economy. The parameters a, b, and k are assumed common to all countries, but the crucial parameter c, which measures the persistence of shocks, differs across countries. This is a convenient way of capturing the idea that differences in institutional arrangements (e.g., the length and timing of wage contracts, or the degree of wage indexation) will affect the persistence of any inflationary shock.

Shocks are assumed to have the same variance in all countries. The reserve country is unconstrained by the pegged-rate regime and chooses its inflation rate using equation (9), with $c = c^*$ (* denotes a reserve country variable):

$$\pi^*_t = ab(k - c^*\varepsilon^*_{t-1}) - [(ab)/(1 + a^2 b)]\eta^*_t \tag{15}$$

The government of a nonreserve country chooses its inflation rate by maximizing equation (14), subject to equations (5), (6), and (15), after observing the realizations of the shock in each country. It chooses the inflation rate

$$\pi_t = ab\left\{k - (1 + h)^{-1}(c\varepsilon_{t-1} + hc^*\varepsilon^*_{t-1}) - g\left[(1 + a^2 b)\eta_t + h\eta^*_t\right]\right\} \tag{16}$$

where $g = \left(1 + a^2 b\right)^{-1}\left(1 + a^2 b + h\right)^{-1}$

The implications of equation (16) are as follows. As h tends to zero, equation (16) tends to equation (9), while as h tends to infinity, it tends to equation (15), and π approaches π^*. For intermediate values of h, we get a weighted average of these two solutions.

Following the same procedure as before, we find that inflation persistence is given by

$$\operatorname{corr}\left(\pi_t, \pi_{t-1}\right) = \left(c + h^2 c^*\right) \div \left\{c^2 + h^2 c^{*2} + g^2 (1+h)^2 \left[\left(1 + a^2 b\right)^2 \left(1 - c^2\right) + h^2 \left(1 - c^{*2}\right)\right]\right\} \tag{17}$$

This expression is effectively a weighted average of inflation persistence under floating rates and inflation persistence in the reserve country, with the weight attached to the latter increasing with the degree of exchange rate constraint (h). It implies that a more constraining exchange rate regime tends to reduce the variance of inflation persistence across countries, because all countries take on the inflation persistence of the reserve currency in proportion to the degree of exchange rate constraint. It does not suggest, as previous authors have speculated, that inflation persistence is necessarily lower in more constraining regimes. The model would only predict this if the exchange rate regime constrains the reserve currency to have low inflation persistence, or if it happened to have low inflation persistence by chance.

The model could be developed in various ways, for example, by allowing more of the parameters to vary across countries. A more realistic version would replace the inflation differential in equation (14) by a term reflecting the accumulated disequilibrium in the real exchange rate. This would reduce to equation (14) only in the special case where the real exchange rate was at its equilibrium value in period t – 1. It seems unlikely, however, that these more complex versions of the model would greatly modify the conclusions reached above. The next section considers the empirical evidence from major OECD countries over the postwar period.

II. Empirical Findings

It may seem a simple matter to estimate inflation persistence across exchange rate regimes, but there is one very significant complication: persistence estimates can easily be biased upward by mean shifts in the inflation process. If we estimate a regression of the form:

$$p_t = a + bp_{t-1} + e_t \tag{18}$$

when the true relationship is:

$$p_t = \alpha_t + \beta p_{t-1} + \varepsilon_t \tag{19}$$

where the t subscript to α indicates that the intercept term varies over time, then b tends to be an upwardly biased estimate of β. This effect is particularly strong if inflation undergoes a significant step shift in mean. The estimated mean obtained from equation (18) [which is equal to $a/(1 - b)$] is then too high for one part of the data set and too low for the other part, so that deviations from the mean appear to be far more persistent than they really are. This effect is sufficient to explain, for example, why much higher estimates of persistence are obtained by estimating equation (18) for OECD countries over the period 1968–99 than over the years 1954–67 (see Bleaney, 1999, for an elaboration of this point with some examples).

The procedure adopted here to deal with this problem is to estimate a modified version of equation (18) whenever there appears to have been a shift in α over the data set. This modification consists of including a dummy variable (D_t) in the regression that takes a value of zero for the first observation and a value of one for the final observation, and shifts from zero to one at a date chosen to maximize the R^2. Thus, the equation estimated is:

$$p_t = a + gD_t + bp_{t-1} + e_t \tag{20}$$

While it may be that the inclusion of D in the regression is erroneous, Monte Carlo estimates show that the downward bias to b that results from the erroneous inclusion of D tends to be much smaller than the corresponding upward bias if D is erroneously omitted (relative to the value of b that results from the correct specification). In 1,000 replications for a data set of 30 observations and either a shift in α by two standard deviations of ε or no shift at all, with a true value of β of 0.50, the mean values of b obtained were as follows: 0.407 when equation (18) was estimated and was the correct specification (i.e., α was constant), 0.401 when equation (20) was estimated and was the correct specification, 0.352 when equation (20) was estimated but equation (18) was the correct specification, and 0.837 when equation (18) was estimated but equation (20) was the correct specification. There is a noticeable downward bias to b even with the correct specification; this bias is, however, irrelevant when comparing b across samples (as here), provided that the bias is approximately equal in each case (which means, in practice, that the comparison involves samples of similar time spans). If the time spans are similar, then the crucial point is that incorrectly omitting D raises b in this case by 0.436 relative to the correct specification, but erroneously including it reduces b by only 0.055. It is thus much safer to risk the latter error than the former.[5]

The analysis is based on annual observations of consumer prices for OECD countries, taken from *International Financial Statistics*.[6] Table 1 gives estimates

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[5]The work of Burdekin and Siklos (1999) does not address the issue of how shifts in α affect estimates of β; they focus on the endogenous choice of break-points for both these parameters.

[6]More frequent observations (monthly or quarterly) are available for most of the period, but it is not clear that it would be advantageous to use them. With higher-frequency observations, measurement error problems would be magnified, since the one-period changes would be smaller. Using annual data also preserves consistency with previous research.

of persistence for sixteen countries for the Bretton Woods period. The first column shows the results of estimating equation (18) over the period 1954–67, and the second column gives estimates from equation (20) over the period 1954–72 with an endogenously chosen shift in the intercept.[7] The estimates of persistence are mostly rather low, with a mean of 0.237 in column (1) and 0.109 in column (2). The difference in means is a consequence of the dummy-variable effect: with an equation (18) specification persistence estimates for 1954–72 are typically greater than 0.5 (Bleaney, 1999, Table 1). It is noticeable that the reserve country (United States) has much higher than average persistence, as also does Canada, which floated its exchange rate from 1950 to 1962.

These findings are not particularly consistent with the theory of the previous section. The variance of persistence across countries appears to be large (although we lack a control group of floating countries), and countries which were not constrained by the exchange rate peg (United States, Canada) seem to have greater persistence, as Alogoskoufis and Smith (1991) suggest.

Table 2 shows the results of a similar exercise for 1984–99. The starting date was chosen to avoid the period of oil shocks and to separate the sample into countries that floated and those that were either members of the exchange rate mechanism (ERM) of the EMS or pegged their exchange rate. The United Kingdom is counted as a floater, although it did participate in the ERM from 1990 to 1992, as is Germany, on the grounds that many observers interpret the EMS as a peg to the German deutsche mark. With eight floating countries and seven nonfloaters, the mean estimate of persistence is almost identical across the two samples. The variance is also lower for the nonfloaters, as predicted by the theory developed above, although not significantly so.[8]

The results for 1984–99 are therefore much more consistent with the theory expounded earlier than those for the Bretton Woods period. It is worth noting, however, that average persistence is significantly higher in the later period than in 1954–72 ($t_{29} = 3.22$, $p < 0.01$).[9] Burdekin and Siklos (1999) also found that there was a significant upward shift in persistence in the 1974–81 period in each of the four countries which they investigated.

Monetary Accommodation and Persistence

It is a basic proposition of international economics that the monetary policy of an individual country is constrained under a system of fixed exchange rates. Since domestic prices need to be kept in line with foreign prices, purely domestic inflation shocks cannot be accommodated by monetary growth in the way in which

[7]The first observation of the dependent variable is not used i.e. when the period is given as 1954–67, the first observation of the dependent variable used is 1955, to allow for the lagged dependent variable. Data earlier than 1954 were not used because of Korean war effects and other disturbances to the inflation process in some countries.

[8]The test for equality of variances yields F(8, 7) = 3.01, compared with a 95% critical value of 4.90.

[9]The *t*-test for the difference of means assumes that both populations are normally distributed with equal variances. The Mann-Whitney U-test for the difference of medians, which assumes only that the two populations have similar distributions, yields a test statistic of 48, which is also significant at the 1% level.

Table 1. Inflation persistence in the Bretton Woods period

	1954–67	1954–72	
	(no breaks)	(one break)	Break at end
Country			
United States	0.476	0.444	1965
Japan	0.390	–0.071	1959
Australia	0.158	0.162	1970
New Zealand	–0.151	0.110	1966
Austria	–0.084	–0.278	1961
Belgium	0.518	0.098	1963
Denmark	0.456	0.039	1961
Germany	0.316	0.300	1970
Ireland	0.229	0.253	1968
Italy	0.513	0.387	1961
Netherlands	0.273	–0.129	1963
Norway	–0.122	–0.331	1969
Sweden	0.067	0.026	1969
Switzerland	0.306	0.298	1968
U.K.	0.212	0.334	1967
Mean (s.d.) of the above	*0.237*	*0.109*	
	(0.226)	*(0.236)*	
Canada	0.687	0.271	1965

Note: Figures given are point estimates of b_1 from the OLS regression $\pi_t = b_0 + b_1 D + b_2\pi_{t-1} + u_t$ where π denotes consumer price inflation; $D = 0$ for 1954 to the break date shown and $D = 1$ for the remainder of the period. The break date is chosen to maximize the R–squared.

they can under floating rates. There is no presumption, however, that monetary policy will *necessarily* be more accommodating under floating rates. Nevertheless, Alogoskoufis and Smith (1991) report evidence of significant monetary accommodation under floating rates, but none under fixed rates, based on an estimate of the world money supply (a weighted average of the United States and the United Kingdom) over more than a century.[10]

This claim needs to be evaluated in light of recent empirical work on central banks' use of interest rate policy. Since the late 1970s policymakers have recognized the potential for expectations to shift the Phillips curve, and, consequently, that monetary policy should react strongly enough to inflation shocks to stabilize expectations (see, for example, OECD, 1977). To achieve this requires essentially that inflation (actual or forecast) have a coefficient of more than one in the nominal interest rate reaction function of the authorities, to ensure that ex ante real interest rates move in the right direction (Svensson, 1999; Taylor, 1999). Clarida and others (1998a) show that the U.S. interest rate

[10]They estimate equation (3) by instrumental variables, using lagged inflation as the instrument. The classification of the Bretton Woods period as a fixed-rate period is questionable, however, since it is not clear that US monetary policy was constrained in any significant way, given that the US dollar was the reserve currency.

Table 2. Inflation persistence and exchange rate regimes 1984–99

	1984–99 (one break)	Break at end
Floating exchange rate		
United States	0.385	1991
Canada	–0.031	1991
Japan	0.554	1991
Australia	0.210	1990
New Zealand	0.318	1987
United Kingdom	0.304	1991
Greece	0.515	1993
Germany	0.482	1995
Mean (s.d.) of the above	*0.342*	
	(0.191)	
EMS or pegged rate		
Austria	0.421	1995
Belgium	0.194	1985
Denmark	0.282	1989
France	0.424	1991
Ireland	0.290	1992
Italy	0.524	1996
Netherlands	0.345	1989
Mean (s.d.) of the above	*0.354*	
	(0.110)	

Notes: Figures given are point estimates of b_1 from the OLS regression $\pi_t = b_0 + b_1 D + b_2 \pi_{t-1} + u_t$ where π denotes consumer price inflation; $D = 0$ for 1984 to the break date shown and $D = 1$ for the remainder of the period. The break date is chosen to maximize the R–squared.

reaction function of the Federal Reserve shifted significantly after 1979, and they estimate the expected inflation coefficient as 0.8 before 1979 and 2.0 afterwards.[11] Clarida and others (1998b) show that post-1979 interest rate reaction functions for other major countries are fairly similar to that for the United States.

These studies treat the interest rate rather than the money supply as the policy variable, with the stock of money adjusting passively to money demand at the given interest rate. This is probably a realistic view of the policy process. Its implication is that, provided that the demand for money function is reasonably stable, monetary accommodation of inflation should be smaller when the interest rate reaction function has a higher inflation coefficient. Thus the evidence from estimated interest rate reaction functions suggests the opposite hypothesis to that of Alogoskoufis and Smith (1991) (hereafter referred to as AS) for the postwar period: monetary accommodation should be found to be stronger in the 1950s and 1960s than in the 1980s and 1990s.

[11]These are the long-run coefficients. The short-run reaction is much less, as there is a high coefficient on the last-period interest rate.

In reality, these two hypotheses are not in direct contradiction and could be combined into a joint hypothesis that the reaction function of the monetary authorities varies over time *and* with the exchange rate regime. For a country which experienced no change in exchange rate regime over time, the former effect should dominate; the latter effect should be evident in a comparison of countries at a *given* date. Since the U.S. dollar was the reserve currency in the Bretton Woods system, the United States is a good approximation to a country with an unchanged exchange rate regime over the postwar period. The estimated model for U.S. monetary growth allows for structural breaks in 1973 (as suggested by AS) and in 1980 (as suggested by Clarida and others). To test for exchange rate regime effects, I use a panel of industrial countries over the period 1983–98. In the empirical implementation, I develop the model slightly by specifying monetary growth as a function of both inflation and monetary growth in the previous year—in other words, I allow for both accommodation of inflation and persistence in monetary growth simultaneously. The reason for this is that financial innovation has shifted the demand function over time, and the gradual spread of financial innovations, as they occur, creates some serial correlation in monetary growth. I find that entering lagged rather than current inflation in the regression yields a better fit (without altering the conclusions in any substantive way), and this permits the use of OLS.

The results for the United States are shown in Table 3. Regression (1) includes shift and multiplicative dummies for a break at the end of the Bretton Woods period and for a break in 1980. According to this regression, monetary persistence is high (the coefficient of lagged monetary growth is around 0.5) but does not appear to vary significantly over the period. Monetary accommodation is estimated to be high initially but to fall substantially after 1980 (the coefficient of lagged inflation is 0.9 for the Bretton Woods period, increases to 1.4 for 1973–79, and falls to 0.3 after 1980). Regression (2) omits the insignificant variables from regression (1), which greatly improves the *t*-statistics. The monetary accommodation coefficient is estimated at 1.05 up to 1979, and then falls to 0.33, with the shift significant at the 5 percent level. These results are more consistent with the story in Clarida and others than with an exchange rate effect.

Given that monetary policy reaction functions appear to have shifted significantly over time, it is important to test for exchange rate regime differences over a period where these reaction functions have been relatively stable. Table 4 reports a similar regression to that reported in Table 3 for a panel of eighteen industrial countries over the period 1984–98.[12] In each year, countries are defined as floating (exchange rate regime dummy = 1) or not floating (exchange rate regime dummy = 0), which includes the membership of the EMS. The classification was based on Ghosh and others (1995), updated to 1998. The coefficients on lagged monetary growth and lagged inflation are allowed to vary according to the exchange rate regime. The results indicate that short-run monetary accommodation was actually *lower* under floating rates, although the signif-

[12]The countries are: Austria, Belgium, Canada, Denmark, Finland, France, Greece, Ireland, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the UK and the US.

[13]The point estimates of long-run accommodation are 0.699 [= 0.633/(1 – 0.094)] for non-floating regimes and 0.601 [= (0.633 – 0.282)/(1 – 0.094 – 0.323)] for floating regimes.

Table 3. Monetary accommodation in the United States 1956–99

Dependent variable: annual percentage growth in money plus quasi-money (Δm_t)

Regression:	(1)	(2)
Independent variables		
Constant	0.902	0.518
	(0.68)	(0.48)
Dummy 1973–99	–8.70	–3.88
	(–1.28)	(–2.32)
Dummy 1980–99	9.07	4.67
	(1.34)	(2.02)
Δm_{t-1}	0.531	0.539
	(2.48)	(4.64)
(Dummy 1973–99)*Δm_{t-1}	0.215	
	(0.48)	
(Dummy 1980–99)*Δm_{t-1}	–0.200	
	(–0.47)	
π_{t-1}	0.912	1.05
	(2.45)	(4.09)
(Dummy 1973–99)* π_{t-1}	0.475	
	(0.77)	
(Dummy 1980–99)* π_{t-1}	–1.06	–0.725
	(–2.02)	(–2.35)
No. of observations	44	44
R–squared	0.610	0.603
Standard error	2.44	2.36
Serial correlation (SC)	0.74	0.23
Functional form (FF)	0.26	0.88
Normality of residuals (N)	0.32	0.46
Heteroscedasticity (H)	1.04	0.93

Note: Figures in parentheses are *t*–statistics. Diagnostic statistics are defined as follows: SC – Lagrange multiplier test for first–order serial correlation (χ^2_1); *FF* – test for functional form based on correlation of residuals with the squared dependent variable (χ^2_1); *N* – Jarque–Bera statistic for normality of residuals (χ^2_2); *H* – test for heteroscedasticity based on correlation between fitted values and squared residuals (χ^2_1).

icantly higher coefficient on the lagged dependent variable means that estimated long-run accommodation is similar.[13]

The results of this section confirm that there was a significant shift in the monetary policy reaction function in the 1970s. On the other hand, there is no evidence that monetary accommodation has been significantly greater in countries with floating exchange rates than in members of the EMS in recent years. This is slightly surprising, since Clarida and others (1998b) find that German interest rates have a large coefficient in estimated interest rate reaction functions of other EMS countries in the post-1979 period, while U.S. interest rates have only a small (although statistically significant) coefficient in the reaction functions of the

Table 4. Monetary accommodation and exchange rate regimes

Eighteen countries 1984–98
Dependent variable: change in the logarithm of money plus quasi–money (Δm_t)

Regression:	(3)
Independent variables:	
Constant	0.0392
	(4.46)
Exchange rate regime dummy	–0.0092
	(–0.70)
Δm_{t-1}	0.094
	(1.18)
(Exchange rate regime dummy)* Δm_{t-1}	0.323
	(2.72)
π_{t-1}	0.633
	(4.40)
(Exchange rate regime dummy)* π_{t-1}	–0.282
	(–1.40)
No. of observations	271
R–squared	0.230
Standard error	0.0638
p–value of test for fixed effects F(17, 248)	0.26

Note: Figures in parentheses are *t*–statistics. Exchange rate regime dummy = 1 for floating rates, = 0 for EMS or pegged rates.

Bundesbank and the Bank of Japan for the same period. It may be that the impact of financial innovation on the money demand function is dominating the results, and effectively blurring the exchange rate regime effects on the dynamics of the money stock.

III. Conclusions

According to the theoretical analysis offered above, the constraints imposed by pegging the exchange rate have the effect of making inflation in all countries reflect the dynamics of inflation in the reserve currency. This only reduces inflation persistence if persistence is particularly low in the reserve country, or if the exchange rate regime constrains the reserve country to have low persistence. It is not clear that either of these last two conditions held under the Bretton Woods system. In the general case, theory predicts that persistence will have a greater cross-country spread under floating than under pegged exchange rates.

Empirical tests of this hypothesis are seriously complicated by variations in the mean inflation rate that have a substantial impact on estimated persistence. There is evidence of significantly increased persistence compared with the 1950s and 1960s in the typical OECD country, but not of higher persistence for floating-rate rather than for EMS countries in the 1980s and 1990s. This is consistent with the Burdekin-Siklos argument that inflation persistence has changed over time for reasons unconnected with the exchange rate regime. The cross-country variance of persistence estimates was surprisingly high in the Bretton Woods period; in the 1980s and 1990s the variance was somewhat lower for EMS than for floating-rate countries, but the sample is small. There is no evidence in recent OECD data that monetary policy is more accommodative under floating rates.

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