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Global Trends in the Choice of Exchange Rate Regime

Michael Bleaney, Mo Tian and Lin Yin

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Michael Bleaney[1], Mo Tian[2] and Lin Yin[3]

[1]School of Economics, University of Nottingham, U.K.

[2]School of Management, Swansea University, U.K.

[3]School of Banking and Finance, University of International Business and Finance, P.R. China

Abstract

The raw data suggest that the global trend towards greater exchange rate flexibility that was evident before 1990 has since stopped. An optimum currency area (OCA) model of exchange rate regime choice is estimated. Four different schemes for classifying exchange rate regime are investigated. Trends in the explanatory variables made little difference to the trend towards greater flexibility before 1990 but have worked against it since, largely because of the reduction in inflation. Underlying preferences are still shifting gradually in the direction of greater flexibility.

Keywords: exchange rate regimes, inflation, openness

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[1]Corresponding author: Professor M F Bleaney, School of Economics, University of Nottingham, Nottingham NG7 2RD, UK. e-mail: michael.bleaney@nottingham.ac.uk. Tel. +44 115 951 5464. Fax +44 115 951 4159.

1. INTRODUCTION

In the 1970s and 1980s, there was a clear trend towards greater exchange rate flexibility: an increasing proportion of countries chose to operate some form of floating exchange rate. As we show below, since 1990 this trend has stopped. The questions that we address here are: does this change represent a switch in preferences for given conditions, or have conditions simply become more favourable to pegs? If the latter, is this likely to continue in the future?

We examine this issue within the framework of an optimum currency area (OCA) model of the choice of exchange rate regime. We find that a float is more likely to be chosen if the inflation rate is above 10%, if the country is rich and large in population, and if it is less open to international trade. Trends in any of these factors will shift regime choices, even if preferences remain the same in the sense that, for given values of the OCA variables, the probability of choosing any given regime remains unchanged. Growing per capita GDP and population should gradually shift choices in the direction of greater flexibility, whereas increasing openness (and the possible sub-division of countries) should operate in the opposite direction. We show that before 1990, OCA factors made very little difference to the trend towards floating; but since 1990, OCA variables have shifted choices decisively towards pegs. By far the most important influence has been the decline in inflation rates in low-income and middle-income countries. This has obscured the fact that there is still an underlying trend in preferences towards greater exchange rate flexibility, although the trend is slower than before 1990. Our results are robust to the choice of exchange rate regime classification scheme. Now that the global disinflation process has largely run its course, observed regime choices are likely to reflect trends in preferences more closely in future.

2. EXCHANGE RATE CLASSIFICATIONS

The appropriate way to classify exchange rate regimes has been the object of a considerable research effort in recent years (see Tavlas *et al*., 2008, for a review). In view of the lack of agreement about the issue, we consider four alternative schemes for which classifications are available for a large sample of countries for all years from 1971 to 2011. The four schemes are those of Shambaugh (2004), Reinhart and Rogoff (2004), Bleaney and Tian (2014), and a modified Shambaugh scheme suggested by Bleaney *et al*. (2015). The modification addresses the fact that the Shambaugh scheme has a rather different approach to devaluations to the others, which to a considerable degree explains its exceptionally low proportion of pegs.

The details of the schemes are[1]:

Shambaugh (2004) [hereafter termed JS]. If the maximum and minimum of the log of the exchange rate against the identified reference currency (the US dollar being the default) do not differ by more than 0.04 over the calendar year, that observation is a peg. Alternatively, if there is a realignment so that the 0.04 threshold is exceeded, the observation is still a peg if the log of the exchange rate is unchanged in eleven months out of twelve. Thus effectively the level of the exchange rate is allowed to vary by ±2%, or alternatively by a realignment in one month and 0% in the remaining eleven months, for a peg to be coded. Note that basket pegs and crawling pegs may well not meet these criteria.

Reinhart and Rogoff (2004) [hereafter termed RR]. Movements of the log of the exchange rate against various reference currencies are analysed. Where available, the exchange rate in the parallel market rather than the official rate is used. If, over a five-year period from years *T–4* to *T*, more than 80% of monthly changes in the log of the exchange rate against any of

[1] The descriptions are taken from Bleaney *et al*. (2015), pp. 3-4 and Table 1.

the reference currencies fall within the range ±0.02, the exchange rate regime in all of the years *T–4* to *T* is classified as some form of peg or band. Alternatively, even if this criterion is not met, if the change in the exchange rate is zero for four months or more, it is classified as a peg for those months. Otherwise it is a float. If the exchange rate moves by more than 40% in a year, that observation is placed in a separate "freely falling" category (these observations are omitted from the analysis in this paper). Thus the scheme focuses on the upper and lower tails of the distribution of monthly exchange rate movements, and specifically the proportion that exceed 2% in absolute value. Note the use of the parallel exchange rate; crawling pegs should meet the criteria for a peg if the crawl is slow enough, but basket pegs may well not do so.

Bleaney and Tian (2014) [hereafter termed BT]. The scheme is based on the root mean square residual (RMSE) from a regression similar to that of Frankel and Wei (1995) for identifying basket pegs. For each calendar year, the change in the log of the official exchange rate against the Swiss franc (the chosen *numéraire* currency) is regressed on the change in the log of the US dollar and of the euro against the Swiss franc. Occasionally, other reference currencies are added.[2] If the RMSE from this regression is less than 0.01, that country-year observation is coded as a peg. If the RMSE is greater than 0.01, twelve new regressions are estimated, each including a dummy variable for a particular month as a test for a realignment. If the F-statistic for the most significant of these dummy variables (April, say) is less than 30, the regime is coded a float. If the F-statistic for April is greater than 30, and the RMSE is less than 0.01, the observation is coded a peg with a realignment; otherwise it is a float. The regression approach should cater for basket pegs (through the regression

[2] See Bleaney and Tian (2014) for details. A similar regression approach to regime classification has been suggested by Benassy-Quéré *et al*. (2006) and Frankel and Wei (2008), but they focus on the estimated coefficients rather than the goodness of fit.

coefficients) or crawls (through the intercept), but errors may arise from the small number of degrees of freedom in each regression.

Bleaney et al. (2015) [hereafter termed BTY]. The scheme replicates the principle of Shambaugh (2004) that for a peg the level of the exchange rate should stay within a ±2% range, after allowing for one possible devaluation, but uses the residuals from a Frankel-Wei regression, as in the case of BT, to capture basket pegs and crawling pegs. The regression period is extended to 24 months (back to January of the previous year) to deal with the problem of lack of degrees of freedom in the BT scheme. The monthly change in the natural logarithm of the exchange rate against the Swiss franc is regressed on the change in the natural logarithm of the US$ and euro rates against the Swiss franc for January of year T-1 to December of year T (with the possible addition of other potential anchor currencies as regressors, as in BT). This regression is repeated 24 times, each with the addition of a dummy for a single month. If the maximum F-statistic for the addition of any monthly dummy is less than 30, the monthly dummies are omitted and the residuals cumulated. Year T is coded as a peg if the maximum cumulated residual minus the minimum cumulated residual < 0.04. If the maximum F-statistic for addition of any monthly dummy is greater than 30, that regression is used in place of the original, and the same criterion of a range of the cumulated residuals of less than 0.04 is applied. Note that, unlike JS, the range permitted for a peg is not reduced to zero in the event of a devaluation.

Figure 1 shows the percentage of pegs recorded by each classification scheme for each year from 1971 to 2011. The JS scheme registers by far the lowest proportion of pegs. The other three schemes record a very similar proportion of pegs up to the late 1980s, but thereafter RR registers a significantly higher proportion than BT or BTY. All four schemes show a shift towards floating up to about 1990, but not since; indeed the JS scheme suggests a mild reversal of this trend between 1991 and 2011. Table 1 shows the estimated coefficient of time for each classification for the two periods. For 1971-90 this coefficient is always negative and highly significant, but the trend is much faster for JS (–1.63% p.a.) than for the other three (–0.63, –0.88 and –0.85% p.a. for RR, BT and BTY respectively). For 1991-2011 JS is again an outlier, showing a highly significant positive trend of 0.68% p.a; for the other three the estimated trend is close to zero and not at all statistically significant.

Figure 1 and Table 1illustrate the question that we wish to address: does this change in trend represent a stabilisation of global preferences since 1990, or can it be explained by changing trends in the independent variables in a model of regime choice? In other words, if we estimate a regime choice model with a time trend, does the time trend display the same shift as Table 1 between the periods before and after 1990? If not, to which variables can this be attributed?

Figure 1. Percentages of pegs identified by year

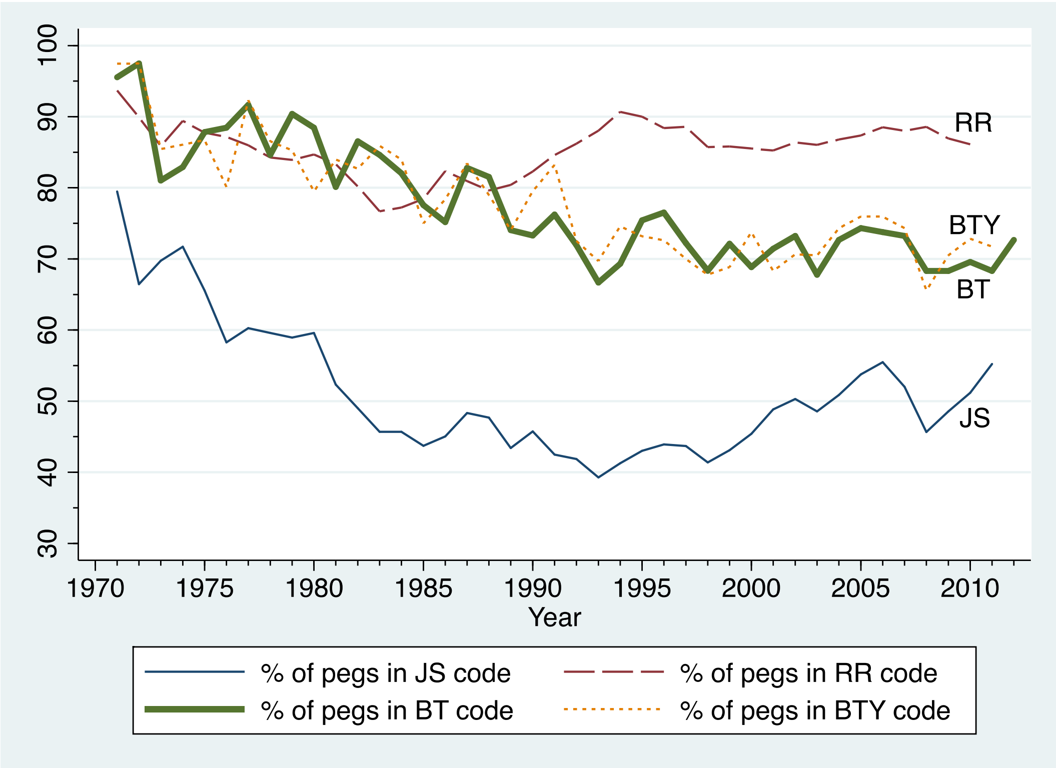

Table 1. Time trends in regime choice, 1971-90 and 1991-2011

Classification scheme:	JS	RR	BT	BTY
1971-90				
Sample size	2966	2128	2932	2934
Time trend	-1.63***	-0.63***	-0.88***	-0.85***
	(-10.71)	(-4.71)	(-7.60)	(-7.58)
1991-2011				
Sample size	3556	2994	3760	3589
Time trend	0.68***	0.021	-0.10	-0.11
	(4.93)	(0.20)	(-0.84)	(-0.91)

Notes. The table shows 100x the coefficient of time (t) in a bivariate regression of Y_{kjt} against t, where Y_{kjt} is a binary regime choice variable (peg=1; float=0) according to classification scheme k in country j in year t. JS: Shambaugh (2004); RR: Reinhart and Rogoff (2004); BT: Bleaney and Tian (2014); BTY: JS24 classification from Bleaney *et al*. (2015). Figures in parentheses are t-statistics. *,**,***: significantly different from zero at the 0.10, 0.05 and 0.01 levels respectively.

3. MODELLING THE CHOICE OF REGIME

The literature on exchange rate regime choice has not yet settled on a definitive model, but the starting point is invariably optimum currency area (OCA) theory, upon which authors generally build to consider a variety of alternative hypotheses. Recent contributions include Bleaney and Francisco (2008), Carmignani *et al*. (2008), von Hagen and Zhou (2007), who provide a comprehensive survey of earlier empirical results, and Levy-Yeyati *et al*. (2010). Variables that typically make their appearance are country size, as measured by GDP, level of development (GDP per capita), openness to international trade (the ratio of exports plus imports to GDP), consumer price inflation, exposure to external shocks (the volatility of the terms of trade), the geographical concentration of trade, financial development (the ratio of M2 to GDP), and more recently liability dollarization (the ratio of foreign liabilities in the banking system to the money stock). Bleaney and Fielding (2002) present a model in which pegging involves choosing the exchange rate before external shocks are observed, but can offer greater anti-inflation credibility, so regime choice involves a trade-off between price stability and output volatility.

In the interests of maintaining the size of the sample (in some cases), and in other cases because they were not statistically significant, not all of these variables are included. Real GDP growth rates, the ratio of foreign exchange reserves to broad money and foreign direct investment as a percentage of GDP were all found to be insignificant. Because it would complicate the analysis of time trends, we also ignore the persistence of exchange rate regimes, as stressed by Bleaney and Francisco (2008) and von Hagen and Zhou (2007), and do not include the lagged regime as a regressor. This means that the statistical significance of persistent explanatory variables is exaggerated by omitted variable bias, but this is not a central concern here.

The explanatory variables that we include are population (as a measure of country size), per capita GDP, openness to international trade and inflation. Instead of the rate of consumer price inflation, we use binary variables for three ranges of the inflation rate: 10 to 20%, 20 to 50%, and above 50%; so the omitted category of inflation is below 10%. This formulation is designed to capture the effects of moderate inflation on regime choice, since the inflation rate coefficient is otherwise liable to be determined by a small minority of observations at the upper end.[3]

The equation that we estimate is a binomial probit for regime *Y* (1=peg; 0=float) in country *j* in year *t* according to classification scheme *k*:[4]

$$Y_{kjt} = \Phi(\mathrm{a}_0 + \mathrm{a}_1\mathrm{INF1020}_{jt} + \mathrm{a}_2\mathrm{INF2050}_{jt} + \mathrm{a}_3\mathrm{INF50}_{jt} + \mathrm{a}_4\mathrm{LPOP}_{jt} + \mathrm{a}_5\mathrm{LPCGDP}_{jt} + \mathrm{a}_6\mathrm{OPEN}_{jt} + a_7\mathrm{t} + \mathrm{u}_{kjt}) \qquad (1)$$

where Φ is the cumulative distribution function of the normal distribution; *INF1020*, *INF2050* and *INF50* are binary variables equal to one for consumer price inflation in the range 10-20% p.a., 20-50% p.a. and greater than 50% p.a. respectively; *LPOP* is the log of population; *LPCGDP* is the log of per capita GDP in 2005 US dollars; and *OPEN* is exports plus imports as a percentage of GDP; the *a*s are parameters to be estimated; and *u* is a random error. The omitted category of inflation is below 10%. One could argue that a simple binary dependent variable is somewhat crude, and that a finer classification scheme for exchange rate regimes should be used. We have not chosen this option because there are several difficulties with it: (1) some categories in a finer classification are underpopulated; (2) their ordering on a scale of increasing flexibility is not always self-evident; and (3) some classification schemes use only a coarse classification.

[3] It is quite common to transform the inflation rate *x* (in %) as 100*x*/(100+*x*); this mitigates but certainly does not eliminate the outlier problem. Country size is also frequently measured by GDP rather than population; in logs GDP is just the sum of per capita GDP and population, so the difference is minor.
[4] Results are similar if we estimate a logit instead of a probit.

4. EMPIRICAL RESULTS

We first estimate equation (1) separately for 1971-90 and 1991-2011, as shown in Table 2. The results are pretty consistent, both across classification schemes and across the two time periods. The coefficients shown are estimated marginal effects, and for the inflation dummy variables they show the difference between a value of one and zero. Inflation is associated with a much higher probability of floating; for 1971-90, compared with inflation below 10%, the estimates show that inflation of between 10% and 20% increases the probability of floating by 5 to 10%, rising to 10% to 20% for inflation between 20% and 50% and to 30% to 40% for inflation over 50%. The one exception is the RR classification, where the coefficient is slightly positive for inflation over 50%, but this is based on a very small sample since RR put most high-inflation observations in a separate "freely falling" category. For 1991-2011 the inflation effect is similar but even a little bit stronger than for 1971-90. The other OCA coefficients (population, per capita GDP and openness) have the expected signs and are mostly significant.

The estimated time trend for 1971-90 is always significantly negative, and is twice as fast for JS (–1.7% p.a.) as for the other three (–0.8% pa.). These figures are very similar to the first row of Table 1, which means that the trend in observed choices matched the trend in preferences, and the net impact of OCA variables on the trend in choices was close to zero. For 1991-2011, the picture is different. Three classification schemes show a significant negative time trend, but in the JS scheme, the estimated time trend is +0.2% p.a., although it is not significantly different from zero. Compared with the lower part of Table 1, the estimated time trend in preferences is about 0.5% p.a. more negative than the time trend in choices, which suggests a trend effect of OCA variables of about +0.5% per annum. A closer analysis reveals that this is entirely the result of reduced inflation, which contributes about +0.4% per annum.[5] Inflation above 20% has become much rarer, as Figure 2 shows.

In order to estimate whether the change in time trend is statistically significant, and since the coefficients for the OCA variables are fairly similar between the two periods, we now estimate a model over the whole period, allowing only the time trend and the intercept term to have different coefficients before and after the end of 1990. In order to allow for the fact that the trend may have been particularly fast in the first few years after the breakdown of the Bretton Woods system, we start the sample at 1976. The results are shown in Table 3. The time coefficient now shows the estimated post-1990 trend in preferences, which is very similar in each case to that shown in Table 2. The coefficient of time multiplied by the 1971-90 dummy shows the estimated difference in trend between the two periods. This coefficient is always negative, indicating a deceleration of the trend in preferences towards floating, and significant at the 5% level in three cases. The estimated shift in trend is particularly large in the case of JS, which as we have previously noted is something of an outlier.

[5] This figure comes from multiplying the time trend in each of the inflation variables over the period since 1991 by its coefficient in the lower panel of Table 2.

Table 2. An optimum currency area model 1971-1990 and 1991-2011

	JS	**RR**	**BT**	**BTY**
From 1971 to 1990				
Observations	1789	1430	1736	1736
Pseudo-R-squared	0.1681	0.1533	0.1385	0.1811
Predicted Prob.	0.4922	0.9145	0.8562	0.8748
time	-0.0171*** (-7.30)	-0.0075*** (-5.73)	-0.0084*** (-5.16)	-0.0075*** (-4.97)
inflation 10 to 20% dummy	-0.0720** (-2.51)	-0.0596*** (-3.14)	-0.0815*** (-3.48)	-0.0770** (-3.36)
inflation 20 to 50% dummy	-0.2048*** (-5.67)	-0.1050*** (-2.85)	-0.1691*** (-4.44)	-0.1586*** (-4.28)
inflation > 50% dummy	-0.4029*** (-11.58)	0.00052 (0.01)	-0.3434*** (-6.16)	-0.3861*** (-6.98)
ln (population)	-0.0643*** (-7.53)	-0.0073 (-1.33)	-0.0365*** (-5.96)	-0.0356*** (-5.62)
ln (real per capita GDP)	-0.1009*** (-12.19)	-0.0194*** (-4.45)	-0.0450*** (-7.95)	-0.0512*** (-9.81)
(exports + imports)/GDP	0.1549*** (4.08)	0.2066*** (7.45)	0.0303 (1.14)	0.0635** (1.96)
From 1991 to 2011				
Observations	2942	2664	2922	2923
Pseudo-R-squared	0.1112	0.1591	0.0938	0.0921
Predicted Prob.	0.4401	0.9144	0.7144	0.7215
time	0.0019 (1.07)	-0.0029** (-3.01)	-0.0065*** (-4.06)	-0.0065*** (-4.08)
inflation 10 to 20% dummy	-0.2289*** (-9.29)	-0.0524** (-2.66)	-0.1058*** (-3.81)	-0.1210*** (-4.38)
inflation 20 to 50% dummy	-0.2741*** (-8.88)	-0.1541*** (-3.70)	-0.2274*** (-5.62)	-0.1852*** (-4.62)
inflation > 50% dummy	-0.4048*** (-16.35)	0.0216 (0.39)	-0.3810*** (-7.83)	-0.3489*** (-7.02)
ln (population)	-0.0531*** (-9.74)	-0.0204*** (-5.46)	-0.0536*** (-10.40)	-0.0533*** (-10.40)
ln (real per capita GDP)	-0.0179*** (-2.84)	-0.0282*** (-8.73)	-0.0233*** (-3.92)	-0.0183*** (-3.11)
(exports + imports)/GDP	0.1205*** (5.17)	0.1715*** (8.50)	0.1086*** (4.61)	0.1068*** (4.66)

Notes. The estimation method is probit, with a binary dependent variable (peg=1; float=0) according to the following classification schemes: JS: Shambaugh (2004); RR: Reinhart and Rogoff (2004); BT: Bleaney and Tian (2014); BTY: JS24 classification from Bleaney *et al*. (2015). Marginal effects at the means of the independent variables are shown. Figures in parentheses are heteroscedasticity-robust *t*-statistics. *,**,***: significantly different from zero at the 0.10, 0.05 and 0.01 levels respectively.

Figure 2. Distribution of inflation rates by year

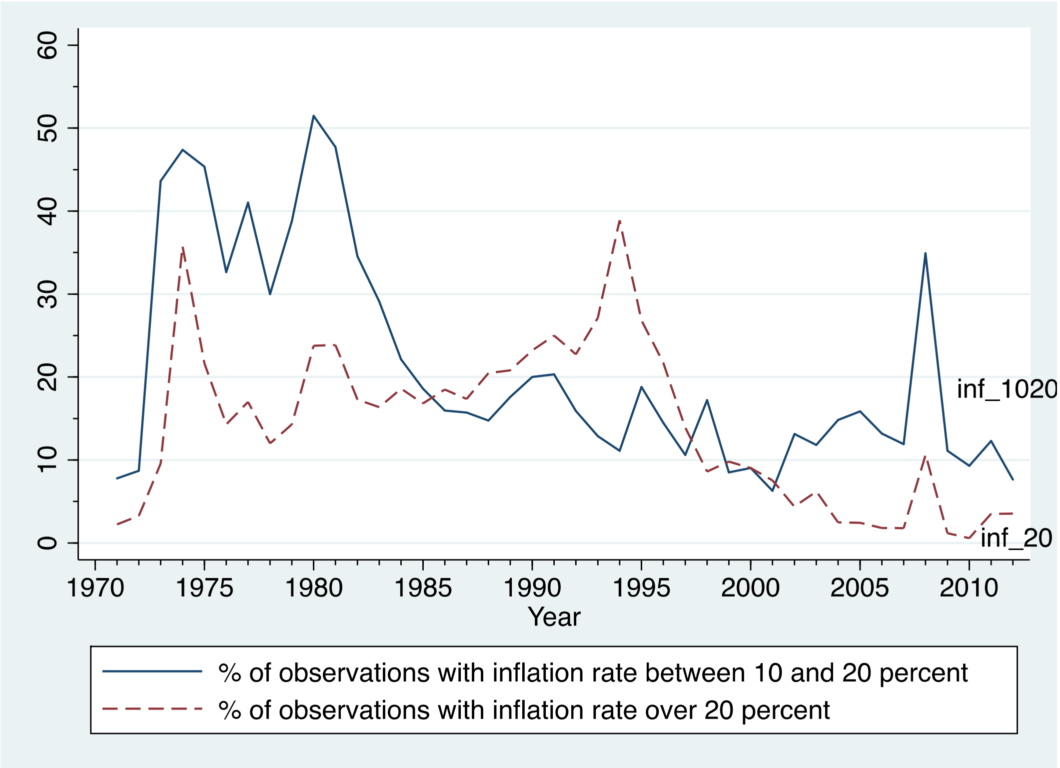

Table 3. Testing for a structural break in the time trend 1976 - 2011

	JS	**RR**	**BT**	**BTY**
Observations	4381	3800	4318	4319
Pseudo-R-squared	0.1215	0.1665	0.1164	0.1212
Predicted Prob.	0.4474	0.9143	0.7628	0.7689
1971-90 dummy	-0.0276 (-0.80)	-0.0785*** (-3.26)	0.00034 (0.01)	0.0460 (1.62)
time * 1971-90 dummy	-0.0200*** (-5.34)	-0.0051** (-2.41)	-0.0102*** (-3.03)	-0.0024 (-0.72)
time (=0 in 1990)	0.0025 (1.40)	-0.0030*** (-3.21)	-0.0058*** (-4.07)	-0.0062*** (-4.37)
inflation 10 to 20% dummy	-0.1833*** (-9.43)	-0.0647*** (-4.27)	-0.1037*** (-4.98)	-0.1179*** (-5.67)
inflation 20 to 50% dummy	-0.2776*** (-12.01)	-0.1483*** (-4.83)	-0.2406*** (-7.74)	-0.2050*** (-6.69)
inflation > 50% dummy	-0.3979*** (-19.92)	0.0096 (0.21)	-0.3723*** (-9.89)	-0.3854*** (-10.44)
ln (population)	-0.0576*** (-12.41)	-0.0180*** (-5.54)	-0.0474*** (-11.49)	-0.0479*** (-11.59)
ln (real per capita GDP)	-0.0393*** (-7.68)	-0.0264*** (-9.95)	-0.0303*** (-6.66)	-0.0320*** (-7.15)
(exports + imports)/GDP	0.1310*** (6.54)	0.1852*** (10.77)	0.0877*** (4.65)	0.1045*** (5.29)

Notes. See notes to Table 2.

5. ROBUSTNESS TESTS

To test the robustness of our results, we estimate a logit model instead of a probit. The results, which are shown in Table 4, are very similar to those in Table 3. With a logit specification, it is also possible to include fixed or random effects. Table 5 shows the estimated marginal effects of a random effects model, from which variables with relatively little time variation (population, per capita GDP and openness) have been omitted. It is not possible to estimate marginal effects from a fixed effects model, but the *t*-statistics in the fixed effects model (not shown) are similar to those for the random effects model. In the random effects model the estimated post-1990 time trend is a bit less negative (or more positive) than in Tables 3 and 4. In the case of BT and BTY this trend is still significantly negative at the 1% level, but the point estimate is about –0.4% per annum rather than –0.6%. For the RR model the post-1990 time trend is no longer significant, although still negative as in Table 3, and for the JS model the positive post-1990 time trend is now statistically significant. The estimated inflation effects are still strong for three of the models, but are surprisingly weak in the RR model.

Table 4. A logit model with a structural break in the time trend 1976-2011

	JS	**RR**	**BT**	**BTY**
Observations	4381	3800	4318	4319
Pseudo-R-squared	0.1217	0.1711	0.1160	0.1203
Predicted Prob.	0.4451	0.9213	0.7674	0.7741
1971-90 dummy	-0.0286 (-0.80)	-0.0621*** (-2.98)	0.0053 (0.18)	0.0457 (1.64)
time * 1971-90 dummy	-0.0208*** (-5.35)	-0.0044** (-2.50)	-0.0095*** (-2.77)	-0.0020 (-0.61)
time (=0 in 1990)	0.0025 (1.39)	-0.0024*** (-3.08)	-0.0058*** (-4.22)	-0.0063*** (-4.57)
inflation 10 to 20% dummy	-0.1840*** (-9.54)	-0.0560*** (-4.15)	-0.1054*** (-4.91)	-0.1188*** (-5.54)
inflation 20 to 50% dummy	-0.2752*** (-12.22)	-0.1332*** (-4.39)	-0.2455*** (-7.54)	-0.2085*** (-6.50)
inflation > 50% dummy	-0.3968*** (-20.26)	0.0115 (0.30)	-0.3809*** (-9.68)	-0.3938*** (-10.31)
ln (population)	-0.0583*** (-11.75)	-0.0143*** (-5.15)	-0.0454*** (-10.71)	-0.0457*** (-10.81)
ln (real per capita GDP)	-0.0394*** (-7.52)	-0.0242*** (-10.92)	-0.0307*** (-6.77)	-0.0319*** (-7.21)
(exports + imports)/GDP	0.1387*** (6.13)	0.1758*** (12.68)	0.0939*** (4.58)	0.1130*** (5.39)

Notes. The estimation method is logit, with a binary dependent variable (peg=1; float=0) according to the following classification schemes: JS: Shambaugh (2004); RR: Reinhart and Rogoff (2004); BT: Bleaney and Tian (2014); BTY: JS24 classification from Bleaney *et al*. (2015). Marginal effects at the means of the independent variables are shown. Figures in parentheses are heteroscedasticity-robust *t*-statistics. *,**,***: significantly different from zero at the 0.10, 0.05 and 0.01 levels respectively.

Table 5. A logit model with random effects 1976-2011

	JS	**RR**	**BT**	**BTY**
Observations	4884	4023	4820	4821
Predicted Prob.	0.5046	0.9986	0.8819	0.9136
1971-90 dummy	-0.0650 (-0.79)	-0.0015 (-0.99)	0.0057 (0.23)	-0.0348 (-1.51)
time * 1971-90 dummy	-0.0547*** (-6.25)	-0.00021 (-1.07)	-0.0090** (-2.48)	-0.0042 (-1.34)
time	0.0136*** (3.06)	-0.000031 (-0.55)	-0.0039*** (-2.61)	-0.0035** (-2.46)
inflation 10 to 20% dummy	-0.1888*** (-4.31)	-0.000030 (-0.07)	-0.0454** (-2.32)	-0.0484** (-2.40)
inflation 20 to 50% dummy	-0.2439*** (-3.78)	-0.0023 (-0.91)	-0.1230*** (-2.77)	-0.0914** (-2.41)
inflation > 50% dummy	-0.3406*** (-3.79)	0.00027 (0.19)	-0.2052*** (-2.72)	-0.2196*** (-3.48)

Notes. See notes to Table 4.

6. CONCLUSIONS

Estimation of an OCA model suggests that the trend in global preferences towards greater flexibility of exchange rates has continued since 1990, despite the fact that floating has not become more common since then. The impact of the trend in preferences on observed choices has been offset in recent years by reduced inflation, which makes pegs more attractive, whereas before 1990 the trend in preferences was fully reflected in observed choices. Three out of the four classification schemes investigated show that this trend has nevertheless decelerated. If the global disinflation process has largely run its course, the recent divergence between trends in preferences and observed choices is likely to disappear.

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