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Global Trends in the Choice of Exchange Rate Regime[1]

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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. The explanatory variables in the model have worked against the trend towards greater flexibility since 1990, largely because of the reduction in inflation.

Keywords: exchange rate regimes, inflation, openness

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1. INTRODUCTION

There have been various empirical studies of the choice of exchange rate regime in recent years, but (partly because of a proliferation of alternative regime classification systems) there is no universally agreed model, nor do we have a clear picture of recent trends in regime choice. Empirical studies of regime choice tend not to investigate time trends, and frequently use time fixed effects to take out the time dimension altogether. This paper has two main aims: to estimate a parsimonious baseline model that is robust to alternative regime classifications and can be used as the stepping-off point for further hypotheses, and to interpret recent trends in regime choice in the light of this model. In particular we show that the trend towards greater flexibility evident before 1990 has stopped, and that this is partly attributable to reductions in inflation having made pegging more attractive.

2. THEORETICAL BACKGROUND

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 Alesina and Wagner (2006), Berdiev *et al*. (2012), Bleaney and Francisco (2008), Breedon *et al*. (2012), Carmignani *et al*. (2008), von Hagen and Zhou (2007), Harms and Hoffmann (2011), Levy-Yeyati *et al*. (2010), Méon and Rizzo (2002) and Rizzo (1998). Optimum currency area theory stresses variables such as country size, as measured by GDP or population, openness to international trade (the ratio of exports plus imports to GDP) and the geographical concentration of trade: small, more open economies with greater geographical concentration of trade are more likely to peg. The inflationary experience of the 1970s led to the development of a different approach that viewed an exchange rate peg as a commitment device that reflects the willingness of the authorities to incur some costs in exchange for price stability. For example 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. From this point of view the inflation rate can be regarded as an indicator of government preferences: governments with a strong desire to insulate output from external shocks and/or a high tolerance for inflation prefer floating rates.[2] The increased frequency of currency crises in a world of ever larger capital flows stimulated the inclusion of variables intended to capture susceptibility to crises, such as the extent of liabilities denominated in foreign currency. A good survey of the empirical literature is provided by von Hagen and Zhou (2007). Political factors have also been investigated, particularly in relation to the discrepancy between self-declared and actual exchange rate regimes that was recognized in the 1990s (Alesina and Wagner, 2006; Carmignani *et al.*, 2008; Méon and Rizzo, 2002).

3. EXCHANGE RATE CLASSIFICATIONS

Any analysis of regime choice requires a system for classifying exchange rate regimes. This is by no means straightforward, and the appropriate way to do it 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

[2] In addition, the algorithms used in some classification schemes may mean that pegs with more than one devaluation in a calendar year, as is likely to occur with rapid inflation, are in any case often classified as floats.

Tian (2014), and a modified Shambaugh scheme suggested by Bleaney *et al*. (2015a). 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 classification of exchange rate regimes in each case is binomial: the regime is either some sort of peg or band, or a type of float (managed or independent). A multinomial classification is also a possibility, but requires more regime boundaries to be identified, and in any case some schemes only provide a binary peg/float classification.

The details of the schemes are[3]:

Shambaugh (2004) [hereafter termed JS]. Each calendar year is analysed separately. 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 of any size 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 the reference currencies fall within the range ±0.02, the exchange rate regime in all of the

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

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]. Each calendar year is analysed separately. 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.[4] 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 coefficients) or crawls (through the intercept), but errors may arise from the small number of degrees of freedom in each regression.

[4] 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.

Bleaney et al. (2015a) [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 floats recorded by each classification scheme for each year from 1971 to 2011 (countries that are members of the Euro Area are counted as pegged from 1999 onwards). The JS scheme registers by far the highest proportion of floats. The other three schemes record a very similar proportion of floats up to the late 1980s, but thereafter RR registers a significantly lower 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 positive 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 negative trend of -0.68% p.a; for the other three the estimated trend is close to zero and not at all statistically significant.

In the context of a model of regime choice, any trend in Table 1 may be broken down into a combination of (a) trends in the explanatory variables, whose impact on regime choice is determined by the coefficients, and (b) a residual trend that we refer to as a trend in preferences, since it represents the trend in the choice that is made for given values of the explanatory variables. One of the aims of the paper is to investigate the role of these two types of trend in determining the observed trend in choices shown in Table 1.

Figure 1. Percentage of floats identified by year

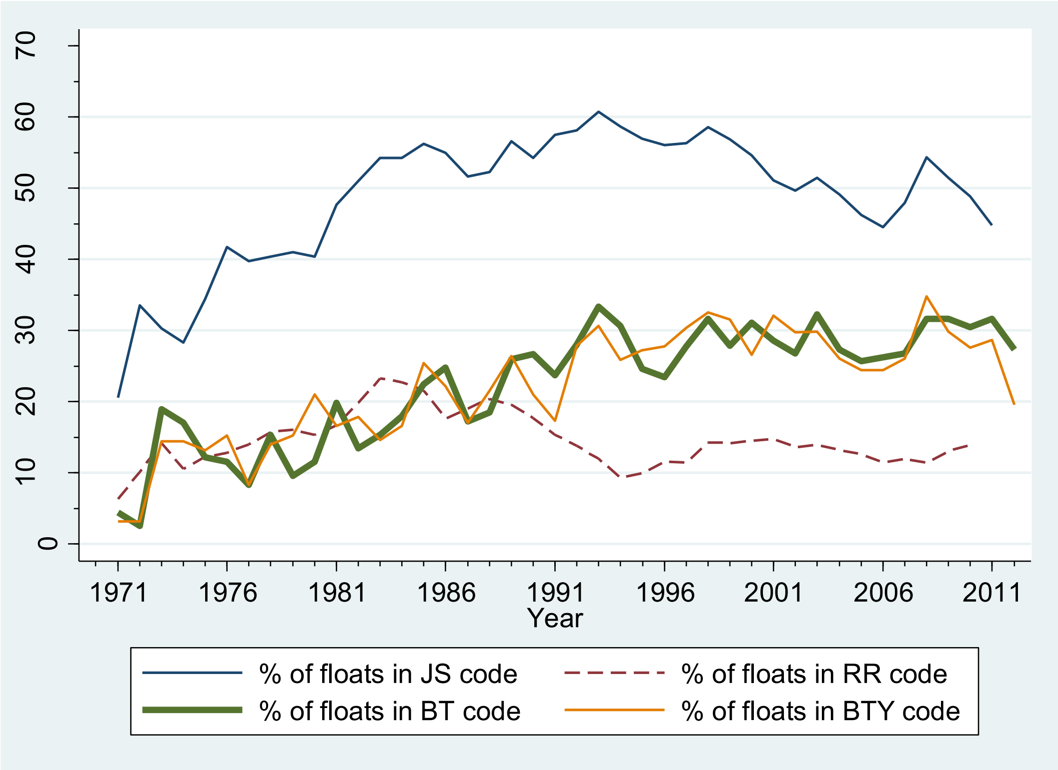

Table 1. Time trends in regime choice, 1971-90 and 1991-2011

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=0; float=1) 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*. (2015a). Figures in parentheses are t-statistics. *,**,***: significantly different from zero at the 0.10, 0.05 and 0.01 levels respectively.

4. MODELLING THE CHOICE OF REGIME

Our model of the probability of floating consists of a time trend plus five other variables: inflation (expected sign: +), the log of population (+), dummy variables for emerging markets and developing countries (+), the ratio of exports plus imports to GDP (-) and the Chinn-Ito (2006) measure of capital account openness (+).[5] The theoretical argument for this last variable is that, with a more open capital account, potential speculative attacks on a currency peg are larger, which may lead the country to prefer some form of float. Other variables, such as GDP growth rates, the ratio of foreign exchange reserves to broad money and foreign direct investment as a percentage of GDP, were considered but were all found to be insignificant. Inflation and the two openness variables are both lagged one year to reduce potential endogeneity. Data sources are given in the Appendix.

Compared to previous research, this model has three main differences: (1) the treatment of country size and level of development; (2) the inclusion of capital account openness; and (3) the specification of the inflation variable.

Country size is usually captured by the log of GDP, which is the sum of the log of population and the log of per capita GDP. We use the log of population, which seems a purer measure of country size than the log of GDP, which is also influenced by the level of development, and we get a better fit using dummies for different levels of development (developed countries, emerging markets, developing countries) than the log of per capita GDP, though the results are quite similar.

We add the Chinn-Ito (2006) measure of capital account openness, which captures potential speculative pressure on an exchange rate peg, so that we expect a more open capital account to be associated with a greater probability of floating. This measure is based on four

[5] We are grateful to a referee for suggesting this variable.

pieces of information: the presence or absence of (a) multiple exchange rates; (b) restrictions on current account transactions; (c) restrictions on capital account transactions; and (d) requirements for the surrender of export proceeds. Because this measure has been developed relatively recently, it has been little used in previous studies (Harms and Hoffmann, 2011, is an exception).

Inflation is a problem because of its highly skewed distribution, with comparatively few extremely high observations. If the inflation rate is entered linearly, these few observations will determine the coefficient no matter what happens at more normal rates of inflation. One approach is to take the logarithm of the inflation rate (Alesina and Wagner, 2006), or to transform it as $p/(1+p)$, where p is the change in the log of the consumer price index (Bleaney and Francisco, 2008; von Hagen and Zhou, 2007), or to include the square of the inflation rate to the regression. These procedures all reduce the problem but fail to eliminate it. A more radical approach that improves the fit, and is used in the working paper version of this article (Bleaney *et al*., 2015b), is to replace the actual inflation rate by a few dummy variables for different ranges, so beyond a certain threshold differences in inflation rates are entirely ignored.[6] Here we achieve the same objective of insulating the coefficient at lower levels of inflation from the influence of outliers by splitting the inflation variable into two: *INF1* is equal to the percentage consumer price inflation rate up to 25%, but is set to 25% if the inflation rate is above 25%, while *INF2* is equal to zero if inflation is less than 25%, and to the actual inflation rate minus 25% for rates above 25%. If the estimated coefficients of these two variables are respectively $b1$ and $b2$, then the estimated effect of inflation (*INF*) is $b1*INF$ if inflation is below 25%, and $25b1+b2*(INF-25)$ if inflation exceeds 25%. If $b1 = b2$, the latter expression also equals $b1*INF$, but it will emerge that the estimates of $b1$ and $b2$ are quite different.

[6] The ranges used were 0 to 10 %, 10 to 20 %, 20 to 50 % and greater than 50 %.

The equation that we estimate, therefore, is a binomial probit for regime *Y* (1=float; 0=peg) in country *j* in year *t* according to classification scheme *k*:[7]

$$Y_{kjt} = \Phi(\mathrm{a}_0 + \mathrm{a}_1\mathrm{INF1}_{jt-1} + \mathrm{a}_2\mathrm{INF2}_{jt-1} + \mathrm{a}_3\mathrm{LPOP}_{jt} + \mathrm{a}_4\mathrm{OPEN}_{jt-1} + \mathrm{a}_5 KA\mathrm{OPEN}_{jt-1} + \mathrm{a}_6\mathrm{EMDUM}_{jt} + \mathrm{a}_7\mathrm{DEVDUM}_{jt} + a_8\mathrm{t} + \mathrm{u}_{kjt}) \quad (1)$$

where Φ is the cumulative distribution function of the normal distribution; *INF1* is the consumer price inflation rate truncated to have a maximum of 25%; *INF2* is inflation minus 25% truncated to have a minimum of zero; *LPOP* is the log of population; *EMDUM* is a dummy variable equal to one for emerging markets and zero otherwise; *DEVDUM* is a dummy variable equal to one for developing countries and zero otherwise; *OPEN* is exports plus imports as a percentage of GDP; is the Chinn-Ito (2006) index of capital account openness; *t* is time (=0 in 2000); the *a*s are parameters to be estimated; and *u* is a random error.

5. EMPIRICAL RESULTS

We first estimate equation (1) separately for 1971-90 and 1991-2011, as shown in Tables 2 and 3. 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 dummy variables they show the difference between a value of one and zero. Inflation up to 25% is always significant at the 1 % level, and the marginal effect of 0.01 implies that an extra 1 % of inflation is associated with an approximately 1 % greater probability of floating. Inflation above 25% has a much smaller coefficient that is generally insignificant. The difference

[7] Results are similar if we estimate a logit instead of a probit.

between the coefficients of *INF1* and *INF2* is always statistically significant, so the data reject a linear specification for inflation.

Country size, as measured by population, always has a positive coefficient, but its marginal effect tends to be greater after 1990 (Table 3) than before (Table 2). In Table 3 it is always significant at the 1 % level, whereas in Table 2 it is only significant at the 5 % level, and in one case (JS) not at all.

Openness to international trade, as captured by the ratio of trade to GDP, always has a significant negative coefficient whose marginal effect lies in the range -0.1 to -0.2, which implies that an extra ten percentage points in the trade ratio is associated with a lower probability of floating of between one and two per cent.

Openness of the capital account, as measured by the Chinn-Ito index, has highly significant positive coefficients after 1990 (Table 3), but much smaller and less significant coefficients in the earlier period. Indeed for the JS classification in Table 2 the estimated coefficient is negative. This difference may reflect greater concern in more recent years about the vulnerability of exchange rate pegs to speculative capital flows in response to major currency crises.

The dummy variables for emerging markets and developed countries have positive coefficients that are significant at the 1 % level in the majority of cases, which indicates that these countries are more likely to float, for given values of the other explanatory variables, than are developing countries. There is quite a lot of variation across classification schemes, with the estimated effect tending to be largest for JS and smallest for RR.

The estimated time trend for 1971-90 is always significantly positive, and is twice as fast for JS (2.0 % p.a.) as for the other three (about 1.0 % pa.). These figures are quite similar to the first row of Table 1, which means that the net effect of any trends in the explanatory variables was close to zero. For 1991-2011, the picture is different. In Table 3 three classification schemes show a significant positive time trend of between 0.2 % and 0.6% p.a., but in the JS scheme, the estimated time trend is -0.2 % p.a., although it is not significantly different from zero.

Table 2. A model of regime choice 1971-1990

Classification:	**JS**	**RR**	**BT**	**BTY**
From 1971 to 1990				
Observations	1615	1310	1577	1596
Pseudo-R-squared	0.1862	0.1617	0.1096	0.1459
Predicted Prob.	0.5936	0.0905	0.1656	0.1594
time	0.0208*** (8.01)	0.0078*** (5.36)	0.0106*** (5.67)	0.0095*** (5.31)
INF1 (lagged)	0.0105*** (5.06)	0.0052*** (5.02)	0.0088*** (6.87)	0.0088*** (7.01)
INF2 (lagged)	0.00327** (2.23)	-0.00074 (-1.57)	-0.00002 (-1.18)	-0.00003 (-1.05)
ln (population)	0.0048 (0.45)	0.0166** (2.29)	0.0199** (2.39)	0.0182** (2.14)
OPEN (lagged)	-0.192*** (-4.82)	-0.207*** (-6.67)	-0.066** (-2.10)	-0.115*** (-3.11)
KAOPEN (lagged)	-0.1193** (-2.41)	0.0390 (1.48)	0.0623* (1.75)	0.1068*** (3.11)
Emerging markets dummy	0.234*** (6.58)	-0.074*** (-5.09)	0.038 (1.05)	0.097** (2.53)
Developed countries dummy	0.407*** (15.9)	0.015 (0.36)	0.186*** (6.69)	0.216*** (7.59)

Notes. The estimation method is probit, with a binary dependent variable (peg=0; float=1) 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.* (2015a). 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. INF1: consumer price inflation (maximum 25%); INF2: consumer price inflation minus 25% (minimum zero); OPEN: (exports + imports)/GDP; KAOPEN: Chinn-Ito capital account openness.

Table 3. A model of regime choice 1991-2011

Classification:	**JS**	**RR**	**BT**	**BTY**
From 1991 to 2011				
Observations	2984	2572	3099	3121
Pseudo-R-squared	0.1487	0.2008	0.1184	0.1254
Predicted Prob.	0.5665	0.0667	0.2854	0.2772
time	-0.0017 (-1.01)	0.0023** (2.52)	0.0061*** (4.13)	0.0047*** (3.28)
INF1 (lagged)	0.0249*** (13.1)	0.0055*** (7.23)	0.0137*** (9.79)	0.0145*** (10.5)
INF2 (lagged)	0.00011 (0.96)	-0.00121** (-2.07)	0.00006 (1.12)	0.00002 (0.93)
ln (population)	0.0234*** (3.70)	0.0165*** (4.54)	0.0280*** (4.89)	0.0284*** (4.98)
OPEN (lagged)	-0.150*** (-5.98)	-0.138*** (-8.48)	-0.128*** (-5.69)	-0.129*** (-5.80)
KAOPEN (lagged)	0.169*** (5.30)	0.064*** (3.91)	0.183*** (6.46)	0.174*** (6.19)
Emerging markets dummy	0.275*** (9.38)	0.060*** (2.60)	0.230*** (6.71)	0.237*** (6.84)
Developed countries dummy	0.103*** (3.75)	0.128*** (5.84)	0.141*** (4.98)	0.146*** (5.13)

Notes. See Notes to Table 2. The estimation method is probit, with a binary dependent variable (peg=0; float=1). Marginal effects are shown.

Table 4 addresses the issue of the apparent structural break around the end of 1990. Equation (1) is estimated over the entire period 1976 to 2011, with the first few years of the post-Bretton Woods era omitted because the shift towards floating was probably particularly fast then. Initially every coefficient was allowed to take a different value up to 1990, but the differences only tended to be statistically significant for three: time and the two dummy variables for emerging markets and developed economies. The structural break test in Table 6 shows that the null hypothesis of no structural break in any coefficient is strongly rejected.

In Table 4 inflation up to 25 %, population, trade openness and capital account openness are all significant at 1 % in every model, with the theoretically expected signs. Emerging markets and developed countries are significantly more likely to float, controlling for other factors, but for emerging markets this is much less true in the 1971-90 period.

We turn now to time trends. The post-1990 trend is shown by the coefficient of TIME in Table 4. This coefficient is significantly positive for three classifications (RR, BT and BTY), but fairly slow (up to 0.5 percentage points p.a.). For all four classifications (but particularly for JS and BT), this is a significant deceleration of the trend towards floating that was evident before 1990 (the estimated shift in the time trend is the coefficient of TIME multiplied by the 1971-90 dummy). This coefficient is always significant at 5 %, but the estimates of the shift vary from 0.46 % p.a. (RR), to 0.88 % p,a. (BTY), 1.46 % p.a. (BT) and 2.57 % p.a. (JS).

Thus the main message of Table 4 is that the shift in preferences towards floating decelerated after 1990, but did not stop entirely.

Table 4. Testing for a structural break in the time trend 1976 – 2011

	JS	**RR**	**BT**	**BTY**
Observations	4326	3647	4441	4448
Pseudo-R-squared	0.1541	0.1978	0.1300	0.1384
Predicted Prob.	0.5735	0.1252	0.2462	0.2421
Structural break	81.3***	61.0***	40.2***	30.4***
1971-90 dummy	0.0164	0.1750***	0.0458	-0.0379
	(0.43)	(5.40)	(1.32)	-1.16)
time	-0.0016	0.0028**	0.0057***	0.0044***
(=0 in 1990)	(-0.96)	(2.98)	(4.30)	(3.33)
time * 1971-90	0.0257***	0.0046**	0.0146***	0.0081**
dummy	(6.51)	(2.46)	(4.35)	(2.48)
INF1 (lagged)	0.0219***	0.0059***	0.0133***	0.0134***
	(15.7)	(9.00)	(12.7)	(13.0)
INF2 (lagged)	0.00023	-0.00103***	0.000015	0.000007
	(1.00)	(-2.66)	(1.21)	(0.567
ln (population)	0.0198***	0.0174***	0.0258***	0.0264***
	(3.67)	(5.08)	(5.37)	(5.56)
OPEN (lagged)	-0.161***	-0.162***	-0.112***	-0.0126***
	(-7.47)	(-10.6)	(-5.98)	(-6.58)
KAOPEN (lagged)	0.114***	0.058***	0.162***	0.166***
	(4.15)	(4.05)	(7.04)	(7.24)
Emerging markets	0.279***	0.063***	0.222***	0.226***
dummy (EM)	(10.0)	(2.67)	(6.78)	(6.86)
Developed cos	0.120***	0.142***	0.138***	0.135***
dummy (DEV)	(4.74)	(6.83)	(5.45)	(5.38)
EM* 1971-90	-0.088	-0.073***	-0.143***	-0.095***
dummy	(-1.47)	(-11.1)	(-5.08)	(-2.84)
DEV * 1971-90	0.237***	-0.068***	0.051	0.097**
dummy	(7.23)	(-9.07)	(1.31)	(2.42)

Notes. See notes to Table 2. The estimation method is probit, with a binary dependent variable (peg=0; float=1). Marginal effects are shown. The structural break test is a test of the joint hypothesis that all four variables that include the 1971-90 dummy have zero coefficients, and is distributed as chi-squared with four degrees of freedom.

Table 1 showed that there was no longer a shift towards floating in observed regime choice after 1990, whereas Table 3 indicates that there was still a significant shift in preferences towards floating (i.e. for given values of the explanatory variables). This difference suggests that the evolution of the explanatory variables since 1990 has operated in favour of pegging, cancelling out the underlying gradual shift in preferences towards floating. In Table 5 we investigate this a bit further.

The main part of Table 5 is a probit for the period 1991-2011, like Table 3, but without *INF2*, which tended to be insignificant. Thus Table 5 assumes that inflation above 25 % has the same effect on regime choice as an inflation rate of exactly 25 %, since *INF1* is truncated at 25 %. The last three rows of Table 5 refer to a modification in which *INF1* is replaced by a detrended version of itself (*RESINF1*); this only changes the coefficient of the time trend (effectively the time trend now includes the effect of the trend in inflation that has been removed). The estimated time trend with *RESINF1* is always less positive (or more negative in the case of JS) than with *INF1*, because of the downward trend in inflation over the period. The last two rows of Table 5 show that the difference is significant at the 1 % level for three out of the four classifications. Indeed when *INF1* is replaced by *RESINF1*, the time trend is very similar to the raw trend in regime choice shown in Table 1, which implies that the downward trend in inflation is responsible for the difference between the trend in preferences and the trend in observed regime choice.

Table 5. A model of regime choice 1991-2011 with inflation truncated at 25 % p.a.

Classification:	**JS**	**RR**	**BT**	**BTY**
From 1991 to 2011				
Observations	2984	2572	3099	3121
Pseudo-R-squared	0.1485	0.1983	0.1176	0.1252
Predicted Prob.	0.5648	0.0689	0.2849	0.2772
time	-0.00180 (-1.04)	0.00236** (2.58)	0.00596*** (4.07)	0.00469*** (3.25)
INF1 (lagged)	0.0252*** (13.5)	0.0051*** (6.81)	0.0140*** (10.1)	0.0146*** (10.6)
ln (population)	0.0236*** (3.72)	0.0171*** (4.63)	0.0282*** (4.92)	0.0285*** (4.99)
OPEN (lagged)	-0.150*** (-5.97)	-0.139*** (-8.38)	-0.128*** (-5.67)	-0.129*** (-5.80)
KAOPEN (lagged)	0.169*** (5.28)	0.067*** (4.00)	0.181*** (6.43)	0.174*** (6.17)
Emerging markets dummy	0.275*** (9.37)	0.060*** (2.57)	0.230*** (6.72)	0.237*** (6.85)
Developed countries dummy	0.104*** (3.76)	0.126*** (5.74)	0.142*** (5.01)	0.146*** (5.14)
Estimated time trend with RESINF1	-0.00885*** (-5.16)	0.00094 (1.05)	0.00205 (1.47)	0.00060 (0.44)
Difference in time trend	-0.00705***	-0.00142	-0.00391***	-0.00408***
t-statistic	-4.08	-0.97	-2.66	-2.83

Notes. See Notes to Table 2. The estimation method is probit, with a binary dependent variable (peg=0; float=1). Marginal effects are shown. The variable RESINF1 is the residuals from a regression of lagged INF1 on a time trend from 1991 to 2011. The last three rows of the table show the estimated time trend when RESINF1 is substituted for lagged INF1 in the regression, the difference between the two estimated time trends, and the *t*-statistic of this difference.

6. ROBUSTNESS TESTS

As stressed by Bleaney and Francisco (2008) and von Hagen and Zhou (2007), exchange rate regimes are persistent. Equation (1) does not include the lagged regime as a regressor, and so the statistical significance of the explanatory variables in equation (1) is exaggerated by this element of pseudo-replication. In Table 6 we test the robustness of the results for 1991-2011 by adding a lagged dependent variable.

The lagged dependent variable is highly significant, with coefficients of 0.849 (JS) and 0.897 (RR), and rather lower ones of 0.521 for BT and 0.556 for BTY, so the lagged regime is a good predictor of the current regime. The coefficients tend to be smaller and less significant than in Table 3, particularly for JS and RR (because the lagged regime explains almost everything), although all the explanatory variables remain significant at the 5 % level with the expected signs for BT and BTY. The coefficients of the time trends are of the same sign as in Table 3, but are no longer statistically significant. Nevertheless Table 6 broadly confirms the significance of the explanatory variables used in the model. Inflation is always significant at the 1 % level up to 25 %, but not significant beyond that, and population, trade openness and capital account openness are all significant at 1 % in the BT and BTY classifications. In the JS classification, they are all significant at 10 %, but with coefficients only slightly smaller than for BT and BTY.

Table 6. Allowing for regime persistence 1991-2011

Classification:	**JS**	**RR**	**BT**	**BTY**
From 1991 to 2011				
Observations	2984	2507	3097	3119
Pseudo-R-squared	0.6549	0.8397	0.3052	0.3405
Predicted Prob.	0.5701	0.0160	0.2626	0.2489
time	-0.0041 (-1.62)	0.0012* (1.91)	0.0024 (1.57)	0.0004 (0.25)
INF1 (lagged)	0.0065*** (2.80)	0.0013*** (3.43)	0.0067*** (4.39)	0.0072*** (4.75)
INF2 (lagged)	0.00001 (0.43)	-0.00066 (-1.57)	0.00003 (0.73)	0.00002 (1.64)
ln (population)	0.0178* (1.90)	-0.0007 (-0.31)	0.0194*** (3.12)	0.0185*** (3.04)
OPEN (lagged)	-0.062** (-2.09)	-0.012 (-1.17)	-0.079*** (-3.41)	-0.080*** (-3.44)
KAOPEN (lagged)	0.078* (1.75)	-0.0026 (-0.29)	0.114*** (3.88)	0.099*** (3.35)
Emerging markets dummy	0.121** (2.16)	0.037 (1.39)	0.140*** (3.84)	0.131*** (3.57)
Developed countries dummy	0.001 (0.03)	0.051*** (2.76)	0.076*** (2.75)	0.076*** (2.75)
Lagged dependent variable	0.849*** (79.6)	0.897** (37.4)	0.521*** (28.2)	0.556*** (30.3)

Notes. See Notes to Table 2. The estimation method is probit, with a binary dependent variable (peg=0; float=1). Marginal effects are shown.

We have chosen a date of the end of 1990 in testing for a structural break in Table 4. In Table 7 we examine the effect of dating it two years earlier, at the end of 1988, or two years later, at the end of 1992. Table 6does not show the full results, but just the estimated time trend in the later period and the time trend multiplied by an early-period dummy. The results are fairly similar to those in Table 4. The estimated time trend in the later period is negative for the JS classification, but significantly positive in five out of six cases for the other three classifications. The time trend in the early period is significantly more positive in seven out of the eight cases.

Table 7. Alternative dates for a structural break

Classification scheme:	JS	RR	BT	BTY
Break date: end-1988				
Time trend in second period	-0.00104 (-0.70)	0.00174** (2.01)	0.00546*** (4.55)	0.00528*** (4.48)
Early period dummy * time trend	0.0323*** (6.70)	0.00766*** (3.34)	0.0147*** (3.59)	0.0100** (2.48)
Break date: end-1992				
Time trend in second period	-0.00327* (-1.71)	0.00301*** (2.86)	0.00427*** (2.86)	0.00177 (1.20)
Early period dummy * time trend	0.0212*** (6.06)	0.00059 (0.35)	0.0110*** (3.85)	0.00705** (2.49)

Note. See notes to Table 4, where the structural break is assumed to occur at the end of 1990. The table shows selected coefficients if the break is assumed t occur (a) at the end of 1988; or (b) at the end of 1992.

7. CONCLUSIONS

In this paper a parsimonious binary model of the choice of exchange rate regimes is estimated, and time trends in regime choice investigated, over the period 1971 to 2011. The probability of floating increases significantly with inflation (but only up to a certain level), population and capital account openness, and decreases significantly with trade openness. Controlling for these factors, the popularity of floating was increasing fairly rapidly up until about 1990, since when the trend has been considerably weaker, although still statistically significant according to three classifications. This deceleration, combined with a general fall in inflation rates that has increased the popularity of pegs, has meant that in terms of raw numbers the pre-1990 trend in the proportion of countries that are classified as floating has stopped.

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Appendix

Appendix Table A1. Data sources

Variable	Data source
JS classification	Jay Shambaugh's website: www.gwu.edu/~iiep/jshambaugh
RR classification	Carmen Reinhart's website: www.carmenreinhart.com
BT classification	www.nottingham.ac.uk/economics/news-events/news/papers/1501.aspx
BTY classification	www.nottingham.ac.uk/economics/news-events/news/papers/1503.aspx
Consumer Price Index	World Development Indicators from World Bank website: http://data.worldbank.org/
Total population	
Trade openness	
Capital openness	The Chinn-Ito Index: http://web.pdx.edu/~ito/Chinn-Ito_website.htm

Appendix Table A2. Sample of Countries (182)

Developed Countries (36)[1]	**Euro Area (18)** Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Latvia, Luxembourg, Malta, Netherlands, Portugal, Slovak Rep., Slovenia, Spain; **Major Advanced Economies (7, 3 of them in Euro Area)** Canada, France, Germany, Italy, Japan, United Kingdom, United States; **Other Advanced Economies (14)** Australia, Czech Rep., Denmark, Hong Kong SAR, Iceland, Israel, Korea, Lithuania, New Zealand, Norway, San Marino, Singapore, Sweden, Switzerland.
Emerging Market Countries (18)[2]	Brazil, Chile, China, Colombia, Egypt, Hungary, India, Indonesia, Malaysia, Mexico, Morocco, Peru, Philippines, Poland, Russia, South Africa, Thailand, Turkey.
Other Developing Countries (128)	**Commonwealth of Independent States (11)** Armenia, Azerbaijan, Belarus, Georgia, Kazakhstan, Kyrgyz Rep., Moldova, Tajikistan, Turkmenistan, Ukraine, Uzbekistan; **Asia (19)** Bangladesh, Bhutan, Brunei, Cambodia, Fiji, Kiribati, Laos, Maldives, Micronesia, Mongolia, Myanmar, Nepal, Papua New Guinea, Samoa, Solomon Islands, Sri Lanka, Tonga, Vanuatu, Vietnam; **Europe (8)** Albania, Bosnia and Herzegovina, Bulgaria, Croatia, Macedonia, Montenegro, Romania, Serbia; **Latin America and the Caribbean (28)** Antigua and Barbuda, Argentina, Aruba, Bahamas, Barbados, Belize, Bolivia, Costa Rica, Dominica, Dominican Rep., El Salvador, Grenada, Guatemala, Guyana, Haiti, Honduras, Jamaica, Netherlands Antilles, Nicaragua, Panama, Paraguay, St. Kitts and Nevis, St. Lucia, St. Vincent and the Grenadines, Suriname, Trinidad and Tobago, Uruguay, Venezuela; **Middle East, North Africa, Afghanistan, and Pakistan (19)** Afghanistan, Algeria, Bahrain, Djibouti, Iran, Iraq, Jordan, Kuwait, Lebanon, Libya, Mauritania, Oman, Pakistan, Qatar, Saudi Arabia, Sudan, Tunisia, United Arab Emirates, Yemen; **Sub-Saharan Africa (43)** Angola, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Cape Verde, Central African Rep., Chad, Comoros, Congo, Cote d'Ivoire, Democratic Rep. of the Congo, Equatorial Guinea, Eritrea, Ethiopia, Gabon, Gambia, Ghana, Guinea, Guinea-Bissau, Kenya, Lesotho, Liberia, Madagascar, Malawi, Mali, Mauritius, Mozambique, Namibia, Niger, Nigeria, Rwanda, Sao Tome and Principe, Senegal, Seychelles, Sierra Leone, Somalia, Swaziland, Tanzania, Togo, Uganda, Zambia.

[1] According to the IMF World Economic Outlook (2015);
[2] According to the MSCI Emerging Markets Indexes (excluded Korea and Czech that are classified as developed countries);