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EXCHANGE RATE VOLATILITY IN INTEGRATING CAPITAL MARKETS

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ABSTRACT

This paper investigates the relationship between international capital liberalization and exchange rate volatility. While the effects of a capital controls liberalization on the transaction volume in the foreign exchange market are theoretically unambiguous, the effects on the volatility of exchange rate can have either sign. On one hand, the liberalization leads to increasing economy-wide and investor-specific uncertainty. On the other hand, the augmented number of participants in the market should reduce exchange rate fluctuations. The uncertainty effects should be dominant in the short run, while the increase in the number of traders in the longer run should make the market thicker and tend to reduce volatility. It is shown that, for a sample of countries which have liberalized capital controls in the last 15 years, structural breaks in the process generating exchange rate volatility have occurred very close to the time when liberalization measures were implemented. The results also suggest an increase in volatility after the structural breakpoint.

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1. Introduction

The structure and openness of the international financial markets has significantly changed in the last two decades. Financial deregulation has led to the internationalization of the financial services industry while the progressive liberalization of international capital movements has allowed, and was followed by, increasing cross–country portfolio diversification. These developments have stimulated the growth of the world foreign exchange markets, the size of which has increased enormously in the last decade. A crucial factor has been the separation between trade motivated transactions and total transactions in foreign exchange. There is empirical evidence that most of the turnover in the foreign exchange markets is generated by financial transactions, a large part of which are of a speculative nature (Arcelli et al. [1990 b]).

The objective of this paper is to investigate the relationship between the removal of capital controls and exchange rate volatility. We first present, in section 2, a model of the foreign exchange market which produces a simple relationship between volume of transactions and exchange rate variability. One of the implications of the model is that there exist changes in the structure of financial markets which generate a simultaneous increase in the transaction volume and in the exchange rate variance. Next, in section 3, we turn to the data. We search for possible structural breaks in the stochastic process generating exchange rates and ask whether these breaks could be related to liberalization measures. The conclusion that we reach is that, for most currencies, changes in the process generating the exchange rate have occurred, and they happened close to the time in which major capital market innovations took place.

In section 4, we explore some implications of the model when different exchange rate regimes are taken into account. We also present some empirical evidence regarding the EMS, suggesting that the progressive strengthening of the European exchange rate agreements may have played a role in reducing liberalization–related volatility with respect to non–EMS countries.

2. The model

In order to investigate the relationship between liberalization and exchange rate volatility, consider a foreign exchange market where J agents trade currency A for currency B[1]. For simplicity we assume that the desired net position in currency A of trader j at time t is a linear function of the form:

$$(1) \qquad Q_{tj} = \alpha\,[e^*_{tj} - e_t]$$

where e_t is the spot exchange rate (defined as units of B currency for one unit of A currency) and e^*_{tj} is the reservation exchange rate for trader j, a function of his/her expectations. Differently from the approach based on one representative agent, we want to account explicitly for differences among participants in the market. Traders' heterogeneity can be introduced in the model by assuming that the information known by different individuals is not identical and, therefore, reservation prices (e^*_{tj}) vary among traders. If $e^*_{tj} > e_t$, currency A is priced below trader j's reservation level, and thus he/she will want to have a positive net position in currency A. Vice versa, if $e^*_{tj} < e_t$, currency A is more appreciated than trader j's reservation level, and he/she will want to have a net positive position in currency B.

In equilibrium, the following must hold:

$$(2) \qquad \sum_{j=1}^{J} Q_{tj} = 0$$

that is, the exchange market clears. Together (1) and (2) imply:

[1] The foreign exchange market is modelled by adapting previous work on financial market volatility and transaction volume by Tauchen and Pitts [1983].

(3) $$\frac{1}{J}\sum_{j=1}^{J} e^{*}_{tj} = e_t$$

i.e. the average reservation exchange rate clears the market. From (3) it is immediate that:

(4) $$\mathrm{Var}(e_t) = \frac{1}{J^2}\,\mathrm{Var}\left[\sum_{j=1}^{J} e^{*}_{tj}\right]$$

the variance of the exchange rate depends on the distribution of the traders' reservation exchange rates. Informational shocks hit the economy every period, at both aggregate and individual levels. Assume that the individuals's reservation prices are distributed as follows:

(5) $$e^{*}_{tj} = \phi_t + \psi_{tj}$$

where:

(6) $$E(\phi_t) = \phi \qquad\qquad \mathrm{Var}(\phi_t) = \sigma^2_{\phi}$$

and

$$E(\psi_{tj}) = \psi \qquad\qquad \mathrm{Var}(\psi_{tj}) = \sigma^2_{\psi}.$$

Accordingly, ϕ_t is an economy wide shock, known by all agents, while ψ_{tj} is idiosyncratic noise. By using (5) and (3), we can write the exchange rate as follows:

$$(7) \qquad e_t = \frac{1}{J}\sum_{j=1}^{J} e^*_{tj} = \phi_t + \frac{1}{J}\sum_{j=1}^{J} \psi_{tj}.$$

If the shocks are assumed to be mutually independent, both across traders and over time, from (7) it follows that

$$(8) \qquad \mathrm{Var}(e_t) \equiv \sigma_e^2 = \sigma_\phi^2 + \frac{\sigma_\psi^2}{J}.$$

Other things equal, an increase in the number of traders (J) tends to reduce the exchange rate volatility. On the other hand, given J, a higher variance of both the aggregate and idiosyncratic informational shocks raises the variance of the exchange rate. In what follows, changes in the distribution of informational shocks and in the market size will be related to the changes in the market regulatory structure.

Consider now the determination of the turnover. The volume of transactions in the market at period t is given by the change in the traders's position between period t–1 and t. Using (7), we can rewrite (1) as:

$$(9) \qquad Q_{tj} = \alpha\left[\psi_{tj} - \frac{1}{J}\sum_{j=1}^{J} \psi_{tj}\right].$$

Therefore, the volume of transactions at t is given by:

$$(10) \qquad V_t = \frac{\alpha}{2}\sum_{j=1}^{J}\left|\left[\psi_{tj} - \frac{1}{J}\sum_{j=1}^{J}\psi_{tj}\right] - \left[\psi_{t-1j} - \frac{1}{J}\sum_{j=1}^{J}\psi_{t-1j}\right]\right|.$$

Note that the V_t is not a function of informational shocks at the aggregate level, which in turn determine e_t. The model thus predicts the possibility of price fluctuations with no corresponding fluctuations in the turnover.

Assuming that ψ_{tj} are iid $\sim$ N(0 ,σ^2_ψ), it can be shown that:

$$\text{(11)} \qquad E(V_t) = a\,\sigma_\psi \sqrt{\frac{J\,(\,J-1\,)}{\pi}}.$$

The expected value of the trading volume is an increasing function of the number of traders in the market. Note also that the average transaction volume is function of the variance σ_ψ reflecting the average dispersion of individual reservation prices, but is not related to the average variability of economic–wide shocks[2].

What does this model tell us about the effects of a capital markets liberalization? This approach highlights the importance of the distinction between short run and long run analysis. In the framework of the model, it is natural to expect an increase in the variance of both the aggregate and the idiosyncratic shocks when the liberalization occurs. The regime change brought about by the elimination of capital controls will augment the amount of information needed by the agents trading in the foreign exchange market. It seem plausible that the degree of uncertainty will be higher, as reflected by larger values of both σ_ϕ and σ_ψ. If, in the short run, the number of traders J does not vary substantially, the increasing uncertainty will produce a positive co–movement of the exchange rate variance and the transaction volume. By (8) and (11), both variables are positively related to σ_ψ. In the short run, therefore, price volatility and transaction volume can be expected to rise together[3].

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[2] The model has been developed by assuming stationarity of the process generating private agents' reservation prices. Therefore, it implies stationarity of the stochastic process for the exchange rate. As long as the main conclusions in this section are concerned, this issue is not crucial. The whole model can be easily recasted in terms of non stationary processes, by re–defining e_t as the exchange rate yield (i.e. $\log[\Delta e_t]$ in terms of the notation in the text) and posing $E(\phi) = E(\psi) = 0$.

[3] International portfolio diversification effects are likely to play a crucial role in the short run increase of transaction volume in the foreign exchange market. The increase in the amount of information needed at the time of the liberalization can be better understood in the framework of the portfolio adjustment which is likely to follow the removal of capital controls.

However, the liberalization also tends to increase the number of potential traders and thus expand the size of the market over time. If this is the case, the correlation between exchange rate variability and transaction volume may reverse sign in the long run. As can be seen from (8) and (11), the transaction volume is a positive function of the number of traders (J), while the exchange rate variance is a negative function of J. As J increases, the markets become "thicker" and this, other things equal, reduces the need for large variations in the exchange rate in order to reach an equilibrium price. In the long run, therefore, the financial liberalization would tend to increase transaction volume and reduce exchange rate fluctuations.

There is empirical evidence that actual liberalization policies have generated considerable increase in the volume of transactions in the foreign exchange market (see Arcelli et al. (1990)). The objective of this paper is to investigate whether they have also generated a short run increase in the volatility of exchange rates, with a possible long run pattern reversal due to the thickening of the market.

Of course, the exchange rate volatility also reflects the attitudes of the domestic monetary authorities towards exchange rate fluctuations. Policy related parameters are included in a modified version of the model presented in Section 4. In this case, as expected, the variance of the exchange rate depends also on the elasticity of the government supply function with respect to deviations of the actual e_t from its desired level. However, the main implications of the model are not affected and the distinction between short and long term consequences of the financial liberalization can still be drawn.

3. Empirical Evidence

A first piece of evidence regarding the volatility of exchange rates over time is provided by Table 1. Using daily data from January 1971 to May 1989, monthly coefficients of variation (i.e. sample standard deviation normalized by the sample mean) of bilateral nominal exchange rates were calculated with respect to three countries which have pursued significant liberalization policies in the period. These countries are as follows:

Australia, where interest rate ceilings were rapidly relaxed at the beginning of the 1980's and most foreign exchange controls were removed in December 1983; Japan, where liberalization policies beginning in 1978 led to the elimination of restrictions on foreign ownership in May 1979 (while the complete set of previous liberalization measures were formally recognized in December 1980); and the UK, where foreign exchange controls were gradually relaxed starting in 1977 and completely eliminated in 1979.

Table 1 shows five–year averages of coefficients of variation. Notice that this measure of volatility increases in correspondence with the liberalization periods. The Japanese data also shows a reduction in the average coefficient of variation toward the end of the sample.

Also, considering the period before and after the liberalization, the average monthly coefficient of variation with respect to the US dollar rises from .56 to 1.32 percent for the Australian dollar, from .81 to 1.28 percent for the Japanese Yen, from .86 to 1.33 percent for the English sterling. The volatility of exchange rates is higher in the post–liberalization period.

The *prima facie* evidence presented in Table 1 suggests that a liberalization of capital movements might lead, at least in the short run, to an increase in the exchange rate volatility. The next step is therefore to test more formally the hypothesis of a relation between capital liberalization and exchange rate volatility.

In the framework of the model presented in section 2, the removal of capital controls leads both to an increase of the transaction volume in the foreign exchange market and to a change in the price variability. The volume effects are theoretically unambiguous. They are also supported by an increasing body of empirical literature.

The sign of the effects on exchange rate volatility, however, is not unambiguous. On one hand, liberalization policies can be seen as leading to higher level of uncertainty, both economy–wide and trader–specific. On the other hand, the raising number of traders improve the thickness of the market. In principle, following a liberalization episode, exchange rate variability might either increase or fall. It could be argued that the uncertainty–related effects are immediate, while the number of traders increases only over time. In this case the exchange rate volatility is expected to rise as soon as the liberalization process starts, while the thickening of the market becomes a relevant factor over the longer run.

If financial deregulation leads to a change in the volatility of the exchange rate, it must be the case that a change in the stochastic process generating the variable occurs in correspondence with the liberalization period.

Therefore, a possible testing strategy consists of estimating the timing of a structural change in the stochastic process generating some measure of exchange rate variability and comparing this result with the timing of the actual liberalization episode. Clearly, the reasons why there might be a structural break in the volatility are numerous, including changes in the exchange rate regime or in the distribution of some fundamentals in the economy. The maintained view in this testing strategy assigns a preminent role to changes in the regulatory structure of capital markets.

Notice that actual liberalization policies have been associated both with flexible exchange rate regimes and with regimes of limited exchange rate flexibility. An example of the second case is given by the EMS experience, where the removal of capital controls has followed a process of progressive strengthening of the target zones regime, producing a reduced exchange rate variability with regards to inter–EMS exchange rates.

On the other hand, the three countries in our sample, Australia, Japan and the UK, witnessed a relatively rapid implementation of liberalization measures in 1983, 1978–80 and 1978–79, respectively, under a regime of floating exchange rates.

Following the methodology adopted by Mankiw et al. [1987], two alternative sets of tests were implemented in order to determine the most likely timing for a regime change in the sample period. The first test endogenously determines a point in time as the most likely switch date from one regime to the other. The second test allows the estimates of the parameters characterizing the two regimes to vary gradually over time from the old to the new values, resulting in an estimate of the adjustment speed between the two.

The second test tries to capture the realistic idea that the move from a regime of capital controls to one of liberalized capital flows is not instantaneous. First, liberalization measures are generally implemented over a period of time (from one to three years in our sample countries), suggesting that there is no single liberalization date. Second, once the financial deregulation takes place, its effects on exchange rate volatility can take some time to occur.

A first measure of variability is the monthly coefficient of variation for daily bilateral exchange rates over the time span June 1973–May 1989.

3.1. A step switching test for the change in regime

Suppose that the process for the coefficient of variation (cv_t) of the exchange rate is as follows:

$$(12) \qquad cv_t = s_1 + \epsilon_t \qquad t = 1,2,\ldots, T_s-1$$

$$cv_t = s_2 + \epsilon_t \qquad t = T_s, T_s+1,\ldots, T$$

where T_s is the switch date, i.e. the first period of the new regime. The objective is to use a maximum likelihood procedure to estimate T_s.

If we assume normal errors, the log likelihood function for the model is:

$$\log L = - T \log(\sigma) + \sum_{t=1}^{T_s-1} \log N \left(\frac{cv_t - s_1}{\sigma} \right) + \sum_{t=T_s}^{T} \log N \left(\frac{cv_t - s_2}{\sigma} \right)$$

where s_1 and s_1 are the means of the coefficient of variation in the old and new regime, respectively, σ^2 is the variance of the error ϵ_t and N (.) is the density function of a standardized normal distribution. As in Mankiw et al. [1987] the maximum likelihood value for T_s is found by computing the maximum likelihood estimates of the three parameters of the model (s_1, s_2 and σ^2) for all possible T_s's and then taking the value of T_s with the maximum likelihood.

Once T_s is estimated, the structural change in the mean of the coefficient of variation at the break date can be tested by using a simple log–likelihood test. The null and alternative hypotheses are:

$$H_0 : \quad s_1 = s_2$$

versus:

$$H_1 : \quad s_1 \neq s_2.$$

The corresponding test statistic is:

$$2 \, (\log NR - \log R) \rightarrow \chi_1^2$$

where NR refers to the non–restricted model and R to the restricted model (i.e. s_1=s_2).

Table 2 reports the result of the step–switching test for the three countries in our sample, with Germany and the USA taken as reference cases. The US dollar bilateral rates with the Canadian dollar, Italian lira, and French franc and German mark will constitute our control group.

To begin with, we focus on the coefficient of variation of each domestic currency *vis a vis* the US dollar: this will provide a first empirical test for the hypothesis that the structural breaks are due to the liberalization. The estimated switch points between regimes are March 1978 for Japan, June 1979 for the UK and March 1983 for Australia. Each of these estimated dates falls well within the corresponding liberalization period.

In all cases, the structural break occurs marginally before the actual implementation of the complete set of liberalization measures. While capital controls are completely removed in May 1979 in Japan, in October 1979 in the UK and in December 1983 in Australia, in each of these countries significant steps towards liberalization are already undertaken in the two year span preceding these dates.

It should also be observed that the likelihood ratio test consistently rejects the equality of the mean coefficient of variation across the two sub–periods, the exchange rate volatility being significantly higher after the structural break.

It could be argued that the structural break might be due to a shift in the volatility of the dollar rather than a country–specific change in regime. The dollar behavior undoubtedly plays a central role in determining the upward trend in the bilateral exchange rate variability beginning at the end of the Seventies. In particular, the change in operating procedures of the U.S. monetary policy in November 1979 (with a move from soft monetary targets to a tight non–borrowed reserves target) led to high interest rate volatility between 1980 and 1982 when the *monetarist experiment* was eventually dropped. The new monetary policy procedures are reflected in the increase in the volatility of the dollar during that period. Nonetheless, if the volatility breaks in our sample were due to the U.S. monetary policy, the breaking points in the bilateral rates with the US dollar should tend to coincide for most currencies, whether or not the country has undertaken liberalization policies. However, switch dates differ considerably not only across Australian dollar, Japanese yen and the English sterling, but also between these and the additional currencies reported in Table 2. In particular, the estimated break for the Australian dollar, Japanese yen and the English sterling occur either before or after the change in U.S. policy (March 1978 for Japan and June 1979 for the U.K. — March 1983 for Australia). Note also that the switch points for Italian lira, French franc and German mark (relative to the dollar) happen at the same date (March 1980). The common break–point is not surprising considering that these currencies were tied together by the March 1979 EMS arrangements. This break date is likely to be the outcome of the increased volatility of the dollar following the change in the U.S. monetary policy at the beginning of 1980.

Consider now the bilateral exchange rate between the currencies of the countries where the liberalization took place. Since they have liberalized at different dates, the break in volatility might reflect the effects of the liberalization in either country. In principle more than one break in the series is possible; more plausibly, the break should occur at the date of the most recent policy change.

The evidence in Table 2 is mixed. On one hand, the Australian late liberalization seems to determine the timing of the shift quite consistently across currencies, as the switch points with respect to the bilateral rates with Japan and the UK occur in February 1985 and September 1984, respectively. This is three to six quarters after the final liberalization measures were implemented in Australia. However, as will be shown in the next section, the estimated adjustment period from a regime of limited capital mobility to one of a free capital movement is three years. This may suggest some delays in the adjustment. Also, the radical policy changes enacted at the beginning of 1985 are associated with a strong depreciation of the Australian dollar[4]. The upsurge in volatility in that year is likely to shift the break forward.

On the other hand, the break point for the yen *vis a vis* the sterling occurs earlier than the beginning of the liberalization process. There are several potential explanations for this early estimated break. First, while capital controls were completely eliminated in the UK only in October 1979, significant liberalization policies were already started in 1977; for example, in that year capital controls on non–residents were eliminated and foreign exchange restrictions on financial intermediaries were relaxed. Thus, the exchange rate volatility might have increased in the 1977–79 period because of expectations that capital controls would soon be completely removed. Second, the changes in monetary control procedures in 1976–77, i.e. the introduction of explicit monetary targets, might have increased both interest and exchange rate volatility. Third, and most important, the

[4] As the Australian economy recovered from the 1982/83 recession, the current account deficits as a percentage of gdp doubled, reaching almost 6%. External and internal imbalance led to marked policy changes starting in 1985 (tightening of both monetary and fiscal policies, as well as the adoption of income policies). The Australian dollar, largely depreciated in the first half of 1985, further weakened in mid–1986.

severe crisis of the pound sterling in 1976, reflected by the temporary high volatility of this currency in that year (see for example Figure 4), have biased the estimate of the break point for the pound/Yen rate towards an early date[5]. The above points can help explain the English currency behavior as captured by the additional tests performed in the remaining sections of this paper.

The breaking dates for the Australian dollar, the yen and the sterling *vis–a–vis* the German mark are 1985, April 1987 and December 1981, respectively. Nonetheless, for both the yen and the sterling the switch is toward a regime of *decreased* variability. In the case of the U.K., this result is not surprising, given the decision by the British Chancellor Lawson to peg (implicitly) the pound to the German mark in 1987: the break date and the fall in volatility in this case captures the informal membership of the U.K in the EMS. On the other hand, the considerable reduction in mark/yen volatility might have reflected the medium–long run consequences of liberalization policies as discussed in section 2; however, the estimated break date is quite early for these effects to have occurred.

One may wonder whether the above results depend on choosing the coefficient of variation as the volatility measure. A new round of tests has thus been performed by using monthly variances of the exchange rate daily growth rates (or daily yields). The corresponding results are reported in Table 3[6].

Switch points for Japan are virtually identical to those in Table 2. The Australian break point of January 1984 *vis a vis* the US dollar now coincides with the actual liberalization date (December 1983). The other Australian break points occur in 1985, in accordance with the results in Table 2. The English sterling pattern is also consistent with

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[5] In order to control for the crisis of the pound and its high volatility in 1976 we run step–switching regressions that exclude that year from the sample. In that case, the break point for the pound/yen rate is estimated to occur in November 1977; that is, right in the middle of the liberlization process for the U.K. and Japan.

[6] Additional tests can been performed by using the variance of the daily exchange rate growth rates. Given the characteristics of the their unconditional distribution, non parametric tests are the most appropriate. Nonetheless, a tentative application of the endogenous switching procedure to the F–test for equality of sub–sample variances has produced results which are strikingly consistent with those in Table 3.

the previous results. The estimated break date is October 1977 with respect both to the U.S. Dollar and to the Japanese Yen. This is well inside the liberalization period even if the high volatility of the pound during the 1976 crisis tends to bias the break date towards a date (end of 1977) that is earlier than the final liberalization measures (1978–79).

One of the disadvantages of the test procedure followed so far is that the switch from one regime to the other is identified with a single point in time. However, the liberalization of capital controls is usually a process over time and, in general, it is sensible to talk about a liberalization period rather than an exact liberalization date.

The test in the next section will directly address the issue of gradualness in the switch from one regime to the other. As a result, an estimate of the speed of adjustment between regimes will be obtained[7].

3.2. A logistic switching test

Assume that at each point in time the parameter s of the process generating CV_t is not constant but is a logistic function of time. The *logistic–switching* model will therefore be as follows:

$$(13) \qquad cv_t = s_t + v_t \qquad t = 1, \ldots\ldots, T$$

$$(14) \qquad s_t = (1-L(t))\, s_1 + L(t)\, s_2$$

where s_1 and s_2 have the same meaning as before, v_t is a normal i.i.d. error corresponding

[7] One way to give a degree of confidence to the point estimates of the regime change date is to compute the posterior odds ratio for alternative dates of regime switch. This ratio allows to find the time region over which it is likely that the regime change has occurred, conditioning on the data. This gives a confidence measure of how likely it is that the liberalization actually occurred in the period around the switching point. This posterior odds ratio is computed as:

$$POR = \exp\left(\log L_t - \log L_{T_s} \right)$$

i.e. the posterior odds ratio is the ratio of the likelihood values for different switch dates.

to ϵ_t in the previous section and, finally,

$$L(t)=\frac{\exp(\alpha+\delta t)}{1+\exp(\alpha+\delta t)}.$$

In each period, the estimated value of s_t will be a convex combination of the parameters values in the two regimes. The inflection point of the logistic, occurring at $t=-(\alpha/\delta)$, will give the date at which both regimes will be equally weighted, i.e. $L(t)=(1-L(t))=1/2$. This point will be reported as switching date between regimes. Notice that considering this rather than other points as switching date is arbitrary and the estimates are not directly comparable with those reported in section 3.1. This is because the logistic switching test imposes some restrictions, such as symmetry in the adjustment process, which are absent from the previous testing procedure.

The rate of change of the parameters value from s_1 to s_2 is captured by δ. Notice, though, that s_2 is reached only asymptotically, since $L(t)\rightarrow 1$ for $t\rightarrow\infty$. Nonetheless, it is useful to present an estimate of the time needed to complete some pre–specified part of the adjustment between the two regimes.

Following Mankiw et al. [1987], consider a fraction 1/2 of the adjustment around the inflection point, i.e. the period in which the weight on the new regime passes from 1/4 to 3/4. Define the dates $L(t(1/4))$ and $L(t(3/4))$ implicitly as $L(t(1/4))=1/4$ and $L(t(3/4))=3/4$. The difference between the second and the first date will be the desired estimate of the time period needed to complete one half of the adjustment around the inflection point. It is easy to show that

$$t(3/4)-t(1/4)=\log(9)/\delta$$

where, as expected, the adjustment time is inversely related to the parameter δ.

The empirical evidence is presented in Table 3 and in Figures 1 through 3. The bilateral Japan/US series is characterized by the shortest adjustment time, 11 months with the first date t(1/4) being in February 1977 and the the second date t(3/4) being December 1979. It can be observed that, while the liberalization process started in 1977 in Japan, most liberalization measures were introduced in 1978–79. In this sense the logistic switching model estimates an adjustment period that is shorter than the actual liberalization process. In the Australia/USA case, t(1/4) corresponds to February 1980 and t(3/4) is in December 1983, with 33 month adjustment. This period covers quite precisely the liberalization process in Australia that started in early 1980 and was completed in 1983.

In the UK/USA case the change between regimes is estimated to take place over a period of 111 months: certainly too long an adjustment period. This result can be explained by observing that the extreme exchange rate variability of the pound in 1976 prevents the test from adequately accounting for the differences across the two regimes[8]. In order to control for the extreme volatility of the pound in 1976, we re–run the logistic switching model by excluding that year from the sample. In this case the adjustment period estimate falls to 62 months, starting in February 1976 (t(1/4)) and ending in March 1980 (t(3/4)). While this second regression still results in a relatively long adjustment process between regimes, it captures the entire liberalization period of the pound in the second half of the 1970's.

In conclusion, the empirical evidence suggests that episodes of capital liberalization are associated with significant structural breaks in the variability of the exchange rate, leading to a systematic volatility increase in the short run. In a first stage, thus, uncertainty–related effects of liberalization measures seem to prevail.

[8] The performance of the logistic switching test depends, among other things, on two crucial features of the series, which are its lenght and the size of the step in the volatility of exchange rates from one regime to the other. In the case of the coefficients of variation for the bilateral rate UK/USA, the high volatility in 1976 reduces the step between the two regimes considerably, stretching the fitted curve as it appears in Figure 3.

4. Liberalization and exchange rate regimes: implication for the EMS

The model in section 2 does not take into explicit consideration the Monetary Authorities. It is obvious that their behavior will significantly affect both the level and volatility of the exchange rate. While there is no simple way of embodying policy rules in the process generating private agents differentiated beliefs, the model can be readily modified to provide some insights about the role played by alternative exchange rate regimes if we assume that the government behavior enters the individual j's reservation price only through the i.i.d. information shocks.

Consider a Central Bank's net supply function in terms of foreign currency of the form:

$$(15) \qquad I_t = \lambda\,(e_t - \bar{e}_t)$$

where $\bar{e}_t$ is some objective in terms of the level of exchange rate and λ is a reaction parameter.

The equilibrium condition (2) will now be as follows

$$(2') \qquad \sum_{j=1}^{J} Q_{tj} = I_t$$

which, together with (1), (2) and (6) will yield

$$(7') \qquad e_t = \frac{\alpha J}{\lambda + \alpha J}\left[\, \phi_t + \frac{1}{J} \sum_{j=1}^{J} \psi_{tj} \right] + \frac{\alpha J}{\lambda + \alpha J}\, \bar{e}_t .$$

Note that in the case of a pure floating exchange rate regime, $\lambda=0$, $I_t=0$ for all t . Expression (7') and (7) coincide.

Suppose now that the monetary authority "leans against the wind" with $\lambda>0$. We report results for three specifications of the policy objective $\bar{e}_t$, namely e_{t-1}, $E(e_t)$ and a constant $\bar{e}$.

When $\bar{e}_t = e_{t-1}$, the government behavior will make the exchange rate level function of the infinite series of past information shocks:

$$(16) \qquad e_t = \sum_{i=0}^{\infty} \left[\frac{\lambda}{\lambda+\alpha J}\right]^i \left[\phi_{t-i} + \frac{1}{J} \sum_{j=1}^{J} \psi_{t-i,j} \right]$$

with variance given by

$$(17) \qquad \mathrm{Var}(e_t) = \frac{(\alpha J)^2}{\lambda+(\alpha J)^2} \left[\sigma_\phi^2 + \frac{\sigma_\psi^2}{J} \right],$$

which is clearly lower than the variance (8) characterizing a free floating regime. Observe that $\mathrm{Var}(e_t)$ approaches zero as $\lambda \rightarrow \infty$.

Alternatively, if $\bar{e}_t$ is defined as the rate level expected by the market, i.e. $E(e_t)$, we can derive e_t and its variance as follows:

$$(18) \qquad e_t = \frac{\alpha J}{\lambda+\alpha J} \left[\phi_t + \frac{1}{J} \sum_{j=1}^{J} \psi_{tj} \right] + \frac{\lambda}{\lambda+\alpha J} \left[\phi + \psi \right]$$

$$(19) \qquad \mathrm{Var}(e_t) = \left(\left[\frac{\alpha J}{\lambda+\alpha J} \right]^2 + \left[\frac{\lambda}{\lambda+\alpha J} \right]^2 \right) \left[\sigma_\phi^2 + \frac{\sigma_\psi^2}{J} \right].$$

Notice that in each period e_t will be a convex combination of the free floating exchange rate level (first term in the right hand side of (I.4)) and the constant expected exchange rate level, given the process generating the information shocks (second term in the RHS of (I.4)). The weight on the second term is proportional to the intervention parameter λ. Again, the exchange rate variance approaches zero as λ goes to infinity.

When, finally, $\bar{e}_t = \bar{e}$ for all t, the variance of the exchange rate will be:

$$(20) \qquad \mathrm{Var}(e_t) = \left[\frac{\alpha J}{\lambda+\alpha J}\right]^2 \left[\sigma_\phi^2 + \frac{\sigma_\psi^2}{J}\right].$$

where the variance is inversely related to λ.

Two interesting issues with respect to the EMS countries can be addressed in the framework of this model.

The first concerns the effects of the removal of capital controls while there is a movement towards a regime of limited exchange rate flexibility.

The model helps interpret the standard rationale for late implementation of liberalization measures in most EMS countries, implementation which has come only after a long and gradual process of consolidation of the exchange rate bands. Within the framework of the model, this process can be interpreted not only in terms of positive values for λ associated with a band for e_t, but also in terms of a steady reduction of the dispersion parameters σ_ϕ and σ_ψ during the shift towards a target zone regime. The uncertainty–related effects of the liberalization are then limited by the increased macroeconomic stability.

Nevertheless, the liberalization might increase potential disturbances in the inter–EMS exchange rates in the short run. In order to insure the stability of exchange rates within the bands, therefore, an even greater degree both of exchange rate intervention (a higher λ) and policy coordination could be called for. In the longer run, however, the increasing thickness in foreign exchange markets (together with greater degree of portfolio diversification) should tend to reduce exchange rate volatility[9].

[9] Given the available sample period, the endogenous regime switching techniques described in Section 3 are not suitable to analize EMS–related regime shifts between European currencies. The consolidation of European target zones has resulted in a negative trend in the inter–EMS exchange rate variability throughout the Eighties. This period could be seen as a long adjustment process toward a regime of reduced volatility, which is not yet represented adequately in the sample. Moreover, extensive

The second question refers to possible effects of the EMS agreements on the bilateral exchange rate volatility with respect to non EMS countries. The issue is whether or not the European Monetary System reduces volatility from external sources, such as liberalization policies pursued by non EMS countries.

Table 5 reports step–switching test results for Germany, France, Italy and the Netherlands with respect to Australia, Japan and the UK. In two cases, Japan and the UK, the volatility decreases in the second period.

The Australian case is different because of the importance of the re–direction of both internal and external policies in 1985. Given the successive depreciations of the Australian dollar during the 1985–86 period, the exchange rate variability increases in the late eighties with respect to the rest of the sample.

The UK–related results clearly reflect the informal English membership in the EMS since 1987. The effects of this membership on exchange rate variability are more apparent at the end of our sample.

Most interestingly, the yen average variability is lower from the beginning of 1981 on in the case of Germany and the Netherlands, from the beginning of 1983 on in the case of Italy and from mid–86 on in the case of France. In the framework of the considerations discussed at the end of Section 2, one could argue that these results reflect the long run effects of capital deregulation. Alternatively, one may suspect that the increasing importance of the inter–EMS ties during the could have played a role in reducing the Yen volatility quite consistently across the EMS countries. However, as shown in Tables 1 and 2, this role for the EMS cannot be detected considering bilateral rates with respect to the dollar (the US dollar related volatility increases in the eighties).

While additional empirical work is necessary in order to verify the hypothesis above, the *prima facie* empirical evidence suggests that volatility–reducing effects of the EMS exchange rate agreements could be detected also with respect to non member countries.

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liberalization policies have been implemented only very recently, precluding the possibility of a thorough assessment of the experience.

5. Conclusions

The growth in the volume of transactions in the foreign exchange markets has risen dramatically in the last decade, to an extent that cannot be explained only by the increase in the volume of trade in goods and services. Most of the turnover in these markets is generated by financial operations, a large part of which are speculative in nature. Moreover, this rise in the volume of transactions has occurred jointly with an increase in the degree of exchange rate volatility. Foreign exchange markets have become more unstable leading to concerns about the "excess volatility" of exchange rates under unrestricted capital mobility.

This paper investigated the sign of the relationship between international capital liberalization, and exchange rate volatility. While the effects of a capital controls liberalization on the transaction volume in the foreign exchange market are theoretically unambiguous, the effects on the volatility of exchange rate can have either sign. On one hand, the liberalization leads to increasing economy–wide and investor–specific uncertainty. On the other hand, the augmented number of participants in the market should reduce exchange rate fluctuations. The uncertainty effects should be dominant in the short run, while the increase in the number of traders in the longer run should make the market thicker and tend to reduce volatility.

In section 3 it has been shown that, for a sample of countries which have liberalized capital controls in the last 15 years, structural breaks in the process generating exchange rate volatility have occurred very close to the time when liberalization measures were implemented. The results also suggest an increase in volatility after the structural break point.

While some *caveats* should be kept in mind in assessing the reported empirical evidence (in particular the possible coexistence of different factors determining the break), the empirical evidence proves to be quite consistent with the analytical framework discussed in Section 2.

What considerations do these results suggest about the effects of capital liberalization in the EMS countries? The implications of the model are twofold.

First, the analytical model helps interpret the standard rationale for late implementation of liberalization measures, after the macroeconomic environment becomes sufficiently stable. This allows to reduce the uncertainty–related effects of the liberalization. On the other hand, short run potential disturbances in the inter–EMS exchange rates can still arise from the removal of capital controls, calling for higher degrees of intervention and policy coordination. Nonetheless, volatility–reducing effects are at work along with the increase of market thickness in the longer run.

Second, the empirical evidence suggests a potential role of the EMS in reducing bilateral volatility of EMS currencies with respect to the Yen, calling for additional empirical research in this area.

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Table 1

MONTHLY COEFFICIENTS OF VARIATION OF BILATERAL EXCHANGE RATE

	AU USA	JA USA	UK USA	
1971-1974	0.46	0.63	0.58	
1975-1979	0.56	0.82	0.86	
1980-1984	0.60	1.35	1.18	
1985-1989	1.32	1.21	1.43	
	JA AU	UK AU	JA UK	
1971-1974	0.70	0.77	0.74	
1975-1979	0.81	0.80	0.98	
1980-1984	1.09	1.01	1.35	
1985-1989	1.53	1.59	1.11	
	CA USA	IT USA	FR USA	GE USA
1971-1974	0.22	0.57	0.81	0.91
1975-1979	0.41	0.77	0.82	0.92
1980-1984	0.47	1.06	1.23	1.17
1985-1989	0.47	1.29	1.35	1.43

AUSTRALIA:	mean CV before liberalization	0.56
	after liberalization	1.32
JAPAN:	mean CV before liberalization	0.81
	after liberalization	1.28
UNITED KINGDOM:	mean CV before liberalization	0.86
	after liberalization	1.33

Table 2
STEP-SWITCHING TEST
MONTHLY COEFFICIENTS OF VARIATION
(Sample 73:6-89:5)

	SWITCH DATE	LR TEST	CV FIRST PERIOD	CV SECOND PERIOD	CV
Australia vs. USA	83:3	34.31	0.53	1.28	0.83
Japan vs. USA	78:3	41.97	0.62	1.31	1.1
UK vs. USA	79:6	28.66	0.82	1.34	1.14
Australia vs Japan	85:2	25.55	0.96	1.75	1.17
Japan vs. UK	76:3	12.08	0.74	1.2	1.12
UK vs. Australia	84:9	23.76	0.96	1.69	1.18
Canada vs USA	76:12	15.78	0.27	0.47	0.43
Italy vs USA	80:3	21.59	0.77	1.26	1.05
France vs USA	80:3	16.75	0.92	1.35	1.17
Germany vs USA	80:3	11.54	1.02	1.37	1.22
Australia vs. Germany	85:3	21.07	1.09	1.79	1.28
Japan vs Germany	81:12	14.99	1.2	0.87	1.05
UK vs. Germany	87:4	10.98	1.01	0.55	0.95

Table 3
STEP-SWITCHING TEST
MONTHLY VARIANCE OF EXCHANGE RATE GROWTH RATES
(Sample 73:6-89:5)

	SWITCH DATE	LR TEST	VARIANCE FIRST PERIOD	VARIANCE SECOND PERIOD	VARIANCE COMPLETE SAMPLE
Australia vs. USA	84:1	34.19	0.002	0.006	0.004
Japan vs. USA	78:3	80.77	0.003	0.006	0.005
UK vs. USA	77:10	58.86	0.003	0.006	0.005
Australia vs Japan	85:3	21.89	0.005	0.008	0.006
Japan vs. UK	77:10	17.11	0.004	0.006	0.005
UK vs. Australia	85:3	25.35	0.005	0.008	0.006
Canada vs USA	76:12	39.34	0.001	0.002	0.0022
Italy vs USA	80:12	51.53	0.003	0.006	0.005
France vs USA	80:12	33.91	0.004	0.006	0.005
Germany vs USA	80:12	23.96	0.005	0.006	0.006
Australia vs. Germany	85:3	19	0.005	0.009	0.006
Japan vs Germany	74:6	26.83	0.008	0.005	0.005
UK vs. Germany	(-)	(-)	(-)	(-)	(-)

Table 4
LOGIT-SWITCHING TEST
MONTHLY COEFFICIENTS OF VARIATION
(Sample 73:6-89:5)

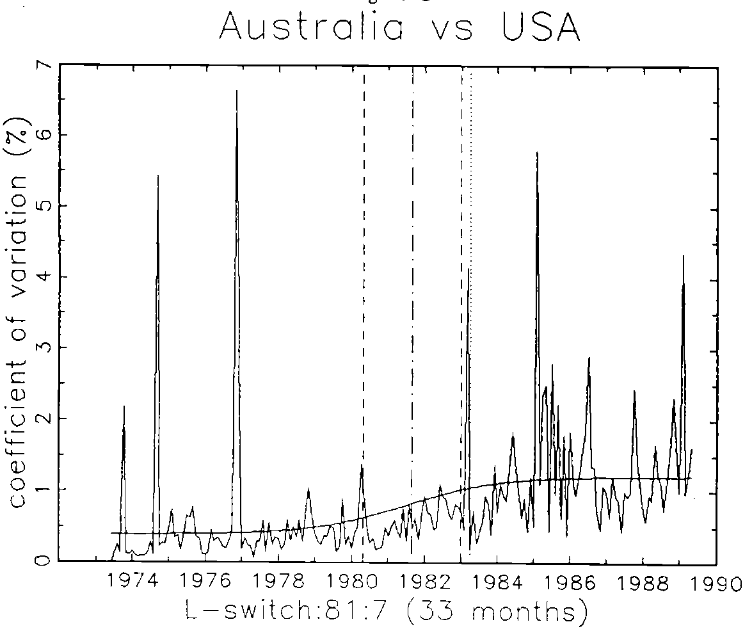
	INFLECTION POINT	MONTHS FOR 1/2 OF THE SWITCH AROUND THE INFLECTION POINT
Australia vs USA	81:7	33
Japan vs. USA	77:6	11
UK vs USA (1)	78:6	62
Australia vs Japan	80:10	40
Japan vs UK	74:3	41
UK vs. Australia	81:4	36

(1) Excluding 1976

Table 5
AUSTRALIA, JAPAN AND THE UK AND THE EMS
(Sample 73:6-89:5)

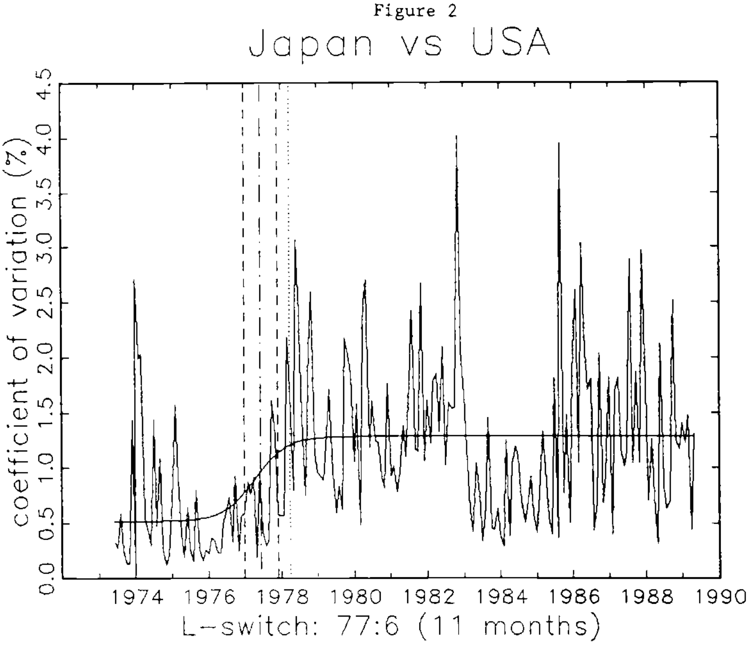
		SWITCH POINT	CV FIRST PERIOD	CV SECOND PERIOD
Australia vs	Germany	85:3	1.09	1.79
	France	85:4	1.04	1.7
	The Netherlands	85:4	1.05	1.78
	Italy	85:4	0.95	1.71
Japan vs	Germany	81:12	1.2	0.87
	France	86:7	1.11	0.74
	The Netherlands	82:1	1.18	0.87
	Italy	83:1	1.12	0.84
United Kingdom vs	Germany	87:5	1.01	0.55
	France	87:6	0.97	0.54
	The Netherlands	87:5	0.97	0.54
	Italy	87:6	0.99	0.52

Figure 1



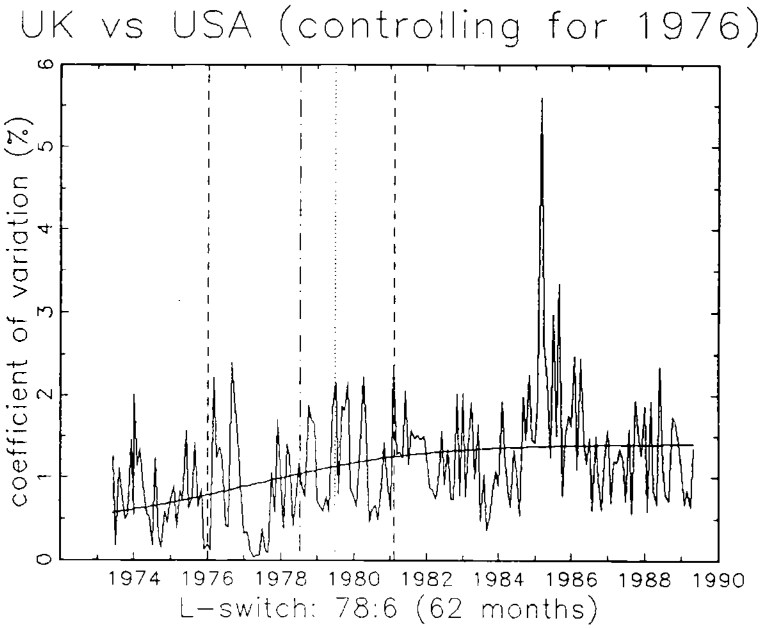
t(1/4) and t(3/4) --------
inflection point _._._._.
step-switching ....

Figure 2



t(1/4) and t(3/4) --------
inflection point _._._._.
step-switching ....

Figure 3

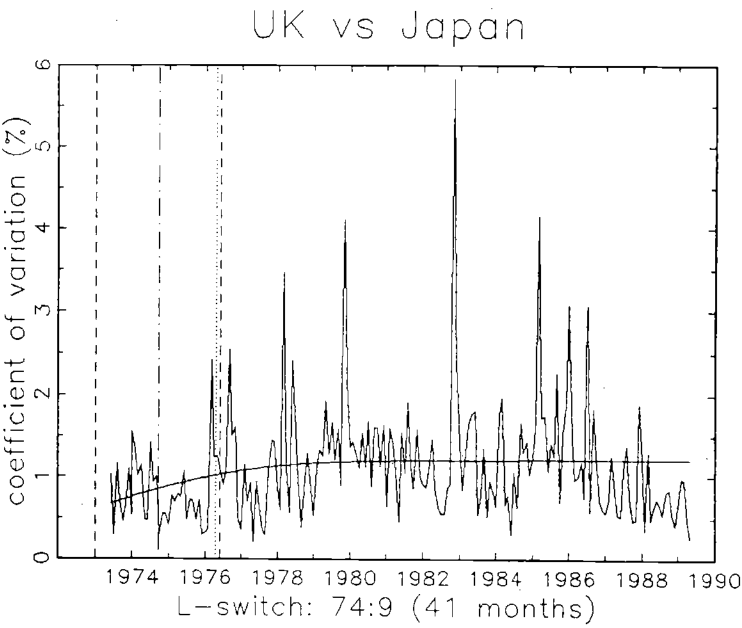


t(1/4) and t(3/4) - - - - - - - -

inflection point _._._._.

step-switching . . . .

Figure 4



t(1/4) and t(3/4) --------
inflection point _._._._.
step-switching ....